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Abstract

In this paper we use a structural VAR model with time-varying parameters and stochastic volatility to investigate whether the Federal Reserve has responded systematically to asset prices and whether this response has changed over time. To recover the systematic component of monetary policy, we interpret the interest rate equation in the VAR as an extended monetary policy rule responding to inflation, the output gap, house prices and stock prices. We find some time variation in the coefficients for house prices and stock prices but fairly stable coefficients over time for inflation and the output gap. Our results indicate that the systematic component of monetary policy in the US, i) attached a positive weight to real house price growth but lowered it prior to the crisis and eventually raised it again, and ii) only episodically took real stock price growth into account.

Keywords: Bayesian VAR, time-varying parameters, monetary policy, house prices, stock market.

JEL classification: C32, E44, E52, E58.

Resumen

En este trabajo utilizamos un modelo VAR estructural con parámetros variables en el tiempo y volatilidad estocástica para investigar si la Reserva Federal ha respondido sistemáticamente a los precios de los activos y si esta respuesta ha cambiado con el tiempo. Para recuperar el componente sistemático de la política monetaria, interpretamos la ecuación de la tasa de interés en el VAR como una regla extendida de política monetaria que responde a la inflación, el *output gap,* los precios de la vivienda y los precios de las acciones. Detectamos variación temporal en los coeficientes de precios de la vivienda y precios de las acciones, mientras que los coeficientes de la inflación y el *output gap* son bastante estables en el tiempo. Nuestros resultados indican que el componente sistemático de la política monetaria en Estados Unidos i) tuvo un peso positivo sobre el crecimiento real de los precios de la vivienda, que disminuyó antes de la crisis y eventualmente volvió a aumentar, y ii) solo tuvo en cuenta el crecimiento real de los precios de las acciones en momentos concretos del tiempo.

Palabras clave: VAR bayesiano, parámetros variables en el tiempo, política monetaria, precios de la vivienda, mercado de valores.

Códigos JEL: C32, E44, E52, E58.

1 Introduction

The length and the severity of the Great Recession generated considerable interest in the evolution of US monetary policy over the period that preceded the recent economic slump. However, while the financial nature of the Great Recession revived the debate on whether monetary policy should respond directly to asset prices (cf. Borio and Lowe (2002), Cecchetti et al. (2002), Bernanke and Gertler (2000, 2001), Christiano et al. (2010), Galí (2014)), less attention has been devoted to the measurement of the actual response of the Federal Reserve to asset prices in recent years.

In this paper we take an empirical approach and evaluate to what extent the Fed reacted to asset prices over the Great Moderation period until the beginning of the Great Recession. In particular, we consider whether stock prices and house prices entered the Fed's reaction function with a positive and significant coefficient. Our key contribution is in providing time-varying estimates of the monetary policy response to asset prices by using a VAR model with time-varying parameters and stochastic volatility (TVP-SV-VAR, henceforth), following Primiceri (2005) and Cogley and Sargent (2005). More specifically, we interpret the interest rate equation in our VAR using five variables (interest rate, inflation rate, the output gap, house prices and stock prices) as an extended monetary policy rule in the spirit of Arias et al. (2015), Belongia and Ireland (2016a,b), Canova and Gambetti (2009) and Primiceri (2005), among others. This set-up allows us to track the systematic response to stock prices and house prices over our sample period, which goes from 1975:Q2 to 2008:Q4. As far as we know, the seminal contributions in this literature (cf. Bernanke and Gertler (2000), and Rigobon and Sack (2003)) and the following extensions are all based on models with constant coefficients. Note that an alternative approach to using a TVP-SV-VAR is to estimate constant coefficient models for either rolling window samples or various sub-samples. However, the TVP-SV-VAR has the advantage of being more flexible as it jointly takes into account the information contained by all the variables in the full sample while at the same time explicitly modeling time variation in both the coefficients and the volatility.

Our main result is that the Fed responded to house prices and stock prices. While the response to stock prices was mild and episodic, the response to house prices was significant, from a statistical and economic point of view. We estimate the coefficient for house price growth to be about one third of the inflation coefficient in the policy rule.

Moreover, we identify non-negligible time variation in the coefficients. The coefficient on stock prices is higher around the end of the 1980s, thus capturing a marked response to the stock market crash of 1987, whereas it is relatively low and stable in the last part of the sample. The coefficient on house price inflation exhibits more pronounced swings: we identify a lower response around the mid 1990s and also in the Pre-Great Recession period. Nevertheless, the coefficient is large, even in the pre-Great Recession period. Finally, the coefficients on inflation and the output gap and the interest rate smoothing term are relatively stable over time, with the partial exception of the mid 1990s.

While we do not find major evidence of time variation in the coefficients for inflation and the output gap, the use of a model with time-varying coefficients and stochastic volatility turns out to be crucial for detecting the Fed's response to house price growth and to stock market returns. In fact, when we shut down time variation in the coefficients or stochastic volatility, the model does not find any response to house price growth. Moreover, the response to stock prices is estimated to be not statistically significant in a model with constant coefficients. Therefore, we conclude that having a model with time-varying coefficients and stochastic volatility is important in order to analyze our research question. Notably, the finding of a significant response to house prices is robust to changing the order of the variables in our VAR.

This paper contributes to two strands of the literature. First, we obviously complement previous studies on the monetary policy response to stock prices. Bernanke and Gertler (2000) estimate Taylor-type rules with a GMM methodology for the US and Japan and find evidence of a very small response, always statistically insignificant and in some cases even negative. Rigobon and Sack (2003) estimate a VAR identified through heteroskedasticity and conclude that the response of monetary policy to stock prices in the US was positive and significant over the period 1985-1999. The same result emerges in

Castelnuovo and Nisticò (2010), in an estimated dynamic stochastic general equilibrium (DSGE) model where monetary policy responds to fluctuations in the stock market, and in Bjørnland and Leitemo (2009), in a VAR identified using a combination of short-run and long-run restrictions. In contrast, a more recent literature argues that the Rigobon and Sack's finding is confined to specific periods (around the 1987 stock market crash in Furlanetto (2011) and more generally around recession periods in Ravn (2012)). While those results rely on various forms of sample splitting, they may highlight some instability in the relationship between monetary policy and stock prices, thus calling for the use of a model with time-varying coefficients.

Interestingly, while several papers study the response of monetary policy to stock prices, the response to house prices is largely unexplored. A noteworthy exception is Finocchiaro and von Heideken (2013) who estimate the house price coefficient in a monetary policy rule and find evidence of a positive and significant response in the US in the context of a DSGE model. Bjørnland and Jacobsen (2013) provide evidence on the (conditional) response of interest rates to shocks originating in the stock market and in the housing sector but do not report the coefficients in the interest rate equation.

We contribute also to a second (and more recent) strand of literature that introduces asset prices into TVP-SV-VAR models. Prieto et al. (2016) use data on several financial variables (including house prices and stock prices) to investigate the time-varying transmission mechanism and the relative importance of various financial shocks. However, they do not consider the systematic component of monetary policy in their analysis. Gali and Gambetti (2015) study the time-varying response of stock prices to monetary policy shocks. While our model is similar, we focus on the opposite relationship, i.e. the response of monetary policy to stock prices.

Finally, we also contribute to the debate on the monetary policy stance in the pre-Great Recession period initiated by Taylor (2007, 2009) who argues that the interest was kept too low for too long prior to the crisis. Belongia and Ireland (2016b) estimate a TVP-SV-VAR model with three variables (a measure of inflation, the interest rate and a measure of real economic activity) and find evidence of a lower response to inflation in recent years, thus supporting the Taylor evidence. Our model can be seen as an extension of their model to include asset prices in the analysis.

The paper proceeds as follows. Section 2 lays out the model and the details of the estimation. Section 3 presents our results and a sensitivity analysis. Section 4 relates our results to the debate on the monetary policy stance in the pre-Great Recession period. Finally, Section 5 concludes.

2 Econometric Model

To study how the Fed responded to asset prices in the pre-Great Recession period, we use the time-varying parameters and stochastic volatility VAR model à la Primiceri (2005)[1] and Cogley and Sargent (2005) with the reduced form representation

$$\boldsymbol{x}_t = \boldsymbol{c}_t + \boldsymbol{B}_{1,t}\boldsymbol{x}_{t-1} + ... + \boldsymbol{B}_{p,t}\boldsymbol{x}_{t-p} + \boldsymbol{u}_t,\ t = 1, ..., T, \tag{2.1}$$

where $\boldsymbol{x}_t$ is a $n \times 1$ vector of endogenous variables, $\boldsymbol{c}_t$ is a $n \times 1$ vector of time-varying coefficients that multiply constant terms, $\boldsymbol{B}_{i,t},\ i = 1, ..., p$ are $n \times n$ matrices of time-varying coefficients and $\boldsymbol{u}_t \sim MVN(0, \boldsymbol{\Omega}_t)$, where $\boldsymbol{A}_t\boldsymbol{\Omega}_t\boldsymbol{A}_t' = \boldsymbol{\Sigma}_t\boldsymbol{\Sigma}_t'$, with $\boldsymbol{\Sigma}_t$ diagonal and $\boldsymbol{A}_t$, the contemporaneous (time-varying) coefficients matrix, lower triangular. In stacked form, the model is equal to:

$$\boldsymbol{x}_t = \boldsymbol{Z}_t'\boldsymbol{B}_t + \boldsymbol{A}_t^{-1}\boldsymbol{\Sigma}_t\boldsymbol{\varepsilon}_t,\ \text{where} \quad \boldsymbol{Z}_t' \equiv \boldsymbol{I}_n \otimes [\boldsymbol{1}, \boldsymbol{x}_{t-1}', ..., \boldsymbol{x}_{t-p}']. \tag{2.2}$$

The time-varying parameters evolve according to $\boldsymbol{B}_t = \boldsymbol{B}_{t-1} + \boldsymbol{\nu}_t,\ \boldsymbol{\alpha}_t = \boldsymbol{\alpha}_{t-1} + \boldsymbol{\zeta}_t$, $\log \boldsymbol{\sigma}_t = \log \boldsymbol{\sigma}_{t-1} + \boldsymbol{\eta}_t$[2]. It is assumed that the innovations in the model are jointly normally distributed with the following variance-covariance matrix:

[1]We follow the updated MCMC procedures suggested by Del Negro and Primiceri (2015). They retain most of the procedures in Primiceri (2005) except that sampling of stochastic volatilities is preceded by sampling of states for mixture component approximations to errors with log chi-square distributions.

[2]To check the validity of these assumptions we run a rolling window constant parameter VAR model. We find random-walk behaviour in the first difference of the coefficients, a result which substantiates the assumptions.

$$V \equiv Var\left(\begin{bmatrix} \boldsymbol{\varepsilon}_t \\ \boldsymbol{\nu}_t \\ \boldsymbol{\zeta}_t \\ \boldsymbol{\eta}_t \end{bmatrix}\right) = \begin{bmatrix} \boldsymbol{I}_n & 0 & 0 & 0 \\ 0 & \boldsymbol{Q} & 0 & 0 \\ 0 & 0 & \boldsymbol{S} & 0 \\ 0 & 0 & 0 & \boldsymbol{W} \end{bmatrix}. \tag{2.3}$$

As a first pass, the structural representation is recovered via a recursive identification scheme. This identification strategy follows the seminal contributions by Christiano et al. (1999) in VAR models with constant coefficients and Primiceri (2005) in VAR models with time-varying coefficients. Notice that our results are robust to different orderings of the variables in the VAR and do not display any particular anomaly that would require a different identification scheme. For instance, we have never encountered counterfactual monetary policy rule coefficients using our recursive identification scheme. This is in line with the results in Arias et al. (2015) which show that structural VARs identified via recursive identification schemes, unlike those identified with restrictions on the sign of impulse responses, do not imply Taylor rule coefficients on inflation and real economic activity with the wrong (negative) sign.

The posterior distributions of $\boldsymbol{B}_t$, $\boldsymbol{Q}$, $\boldsymbol{A}_t$, $\boldsymbol{S}$ and $\boldsymbol{W}$ are obtained via Gibbs sampling[3] with standard prior assumptions as in Primiceri (2005)

$$\boldsymbol{B}_0 \sim N\left(\widehat{\boldsymbol{B}}_{OLS}, 4 \cdot V\left(\widehat{\boldsymbol{B}}_{OLS}\right)\right), \tag{2.4}$$

$$\boldsymbol{Q} \sim IW\left(k_Q^2 \cdot \tau \cdot V\left(\widehat{\boldsymbol{B}}_{OLS}\right), \tau\right), \tag{2.5}$$

$$\boldsymbol{A}_0 \sim N\left(\widehat{\boldsymbol{A}}_{OLS}, 4 \cdot V\left(\widehat{\boldsymbol{A}}_{OLS}\right)\right), \tag{2.6}$$

$$\boldsymbol{S}_m \sim IW\left(k_S^2 \cdot (m+1) \cdot V\left(\widehat{\boldsymbol{A}}_{m,OLS}\right), (m+1)\right), \tag{2.7}$$

$$\log \boldsymbol{\sigma}_0 \sim N\left(\log \widehat{\boldsymbol{\sigma}}_{OLS}, 4 \cdot \boldsymbol{I}_n\right) \tag{2.8}$$

$$\boldsymbol{W} \sim IW\left(k_W^2 \cdot (n+1) \cdot \boldsymbol{I}_n, (n+1)\right), \tag{2.9}$$

[3] The total number of Gibbs sampling iterations is set to 150,000 with a burn-in of 100,000 draws and convergence is checked by means of rolling variances plots. We keep the remaining 50,000 draws and use every 100th for inference. The results are basically identical if, more conservatively, we kept every 20th draw instead. In that case, if anything, we find that the coefficient on S&P 500 growth is significant with a magnitude of 0.02 not only around the 1987 financial crisis but also in 2007:Q3, i.e., prior to the onset of the Great Recession. Also, results are unaffected if we do or we do not truncate the autoregressive matrices to yield stationary draws. In all exercises, the number of stationary draws is always above 2/3.

where $m = 1, ..., n-1$, $\widehat{\boldsymbol{A}}_{m,OLS}$ is the m-th row of $\widehat{\boldsymbol{A}}_{OLS}$ and τ is the size of the training sample on which a time invariant VAR is estimated by OLS in order to calibrate the prior distributions described above. Following Primiceri (2005), we use the first 10 years of data as a training sample to calibrate the priors for estimation over the actual sample period, which starts in 1985:Q3. The selection of the hyperparameters also follows Primiceri (2005) in choosing $k_Q = 0.01$ and $k_S = 0.1$, with the sole exception of allowing for more time variation in the stochastic volatility by setting k_W to 1 as opposed to 0.01 and by tuning the prior variance of $\log \boldsymbol{\sigma}_t$ to $4\boldsymbol{I}_n$ instead of $\boldsymbol{I}_n$. With these choices of hyperparameters, the priors are diffuse and uninformative and, in fact, the prior for the stochastic volatility of the model described in (2.8) is de facto flat. This choice of priors is conservative for the question we address in the sense that it does not restrict the amount of potential time variation in the volatility of the model and, thus, does not artificially blow up the time variation in the policy coefficients.

We consider quarterly data from 1975:Q2 to 2008:Q4. In particular, the vector $\boldsymbol{x}_t = [\Pi_t \quad \widetilde{Y}_t \quad \Delta H_t \quad \Delta S\&P^{500}_t \quad FFR_t]'$ consists of Π_t, year-over-year percentage changes in the deflator for personal consumption expenditures (excluding food and energy), $\widetilde{Y}_t$, the output gap measured as the percentage-point difference between actual real GDP and the US Congressional Budget Office estimate of real potential GDP, ΔH_t, the percentage growth of the real Freddie Mac House price index, $\Delta S\&P^{500}_t$, the percentage growth of the real S&P 500 index, and FFR_t, the federal funds rate. Asset prices are deflated by core PCE. All raw series are drawn from the FRED database.

A lag length of $p = 1$ is suggested by the BIC criterion obtained from OLS estimation of the constant parameters version of our model. This lag order has the fortunate by-product of facilitating the comparison with the macroeconomic literature on Taylor rules with interest rate smoothing.

The systematic component of monetary policy is recovered from the structural representation of our model

$$\boldsymbol{A}_t \boldsymbol{x}_t = \boldsymbol{A}_t \boldsymbol{c}_t + \boldsymbol{A}_t \boldsymbol{B}_{1,t} \boldsymbol{x}_{t-1} + \boldsymbol{\Sigma}_t \boldsymbol{\varepsilon}_t \tag{2.10}$$

or, equivalently, and omitting constant terms

$$\underbrace{\begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ a_{21,t} & 1 & 0 & 0 & 0 \\ a_{31,t} & a_{32,t} & 1 & 0 & 0 \\ a_{41,t} & a_{42,t} & a_{43,t} & 1 & 0 \\ a_{51,t} & a_{52,t} & a_{53,t} & a_{54,t} & 1 \end{bmatrix}}_{\boldsymbol{A}_t} \underbrace{\begin{bmatrix} \Pi_t \\ \widetilde{Y}_t \\ \Delta H_t \\ \Delta SP_t \\ R_t \end{bmatrix}}_{\boldsymbol{x}_t}$$

$$= \underbrace{\begin{bmatrix} ab^1_{11,t} & ab^1_{12,t} & ab^1_{13,t} & ab^1_{14,t} & ab^1_{15,t} \\ ab^1_{21,t} & ab^1_{22,t} & ab^1_{23,t} & ab^1_{24,t} & ab^1_{25,t} \\ ab^1_{31,t} & ab^1_{32,t} & ab^1_{33,t} & ab^1_{34,t} & ab^1_{35,t} \\ ab^1_{41,t} & ab^1_{42,t} & ab^1_{43,t} & ab^1_{44,t} & ab^1_{45,t} \\ ab^1_{51,t} & ab^1_{52,t} & ab^1_{53,t} & ab^1_{54,t} & ab^1_{55,t} \end{bmatrix}}_{\boldsymbol{A}_t \boldsymbol{B}_{1,t}} \underbrace{\begin{bmatrix} \Pi_{t-1} \\ \widetilde{Y}_{t-1} \\ \Delta H_{t-1} \\ \Delta SP_{t-1} \\ R_{t-1} \end{bmatrix}}_{\boldsymbol{x}_{t-1}} + \underbrace{\begin{bmatrix} \sigma_{1,t} & 0 & 0 & 0 & 0 \\ 0 & \sigma_{2,t} & 0 & 0 & 0 \\ 0 & 0 & \sigma_{3,t} & 0 & 0 \\ 0 & 0 & 0 & \sigma_{4,t} & 0 \\ 0 & 0 & 0 & 0 & \sigma_{5,t} \end{bmatrix}}_{\boldsymbol{\Sigma}_t} \underbrace{\begin{bmatrix} \varepsilon_{\pi,t} \\ \varepsilon_{Y,t} \\ \varepsilon_{H,t} \\ \varepsilon_{SP,t} \\ \varepsilon_{R,t} \end{bmatrix}}_{\boldsymbol{\varepsilon}_t}. \tag{2.11}$$

Looking at the fifth row of (2.11), we have

$$\begin{aligned} & a_{51,t}\Pi_t + a_{52,t}\widetilde{Y}_t + a_{53,t}\Delta H_t + a_{54,t}\Delta S\&P^{500}_t + R_t \\ & \qquad = ab^1_{51,t}\Pi_{t-1} + ab^1_{52,t}\widetilde{Y}_{t-1} + ab^1_{53,t}\Delta H_{t-1} + ab^1_{54,t}\Delta S\&P^{500}_{t-1} + ab^1_{55,t}R_{t-1} + \sigma_{5,t}\varepsilon^R_t. \end{aligned} \tag{2.12}$$

Bringing R_t over to the left-hand side yields

$$\begin{aligned} R_t = & -a_{51,t}\Pi_t - a_{52,t}\widetilde{Y}_t - a_{53,t}\Delta H_t - a_{54,t}\Delta S\&P^{500}_t \\ & + ab^1_{51,t}\Pi_{t-1} + ab^1_{52,t}\widetilde{Y}_{t-1} + ab^1_{53,t}\Delta H_{t-1} + ab^1_{54,t}\Delta S\&P^{500}_{t-1} + ab^1_{55,t}R_{t-1} \\ & + \sigma_{5,t}\varepsilon^R_t, \end{aligned} \tag{2.13}$$

where the coefficient $ab^1_{55,t}$ captures the degree of interest rate smoothing.

We will focus on the time evolution of the sum of the coefficients on the contemporaneous and lagged variables (e.g. $-a_{51,t}+ab^1_{51,t}$ for Π_t) but we will also present results for the long-run coefficients (obtained by dividing the sum of coefficients by $(1-ab^1_{55,t})$) in order to represent the response of the interest rate to a permanent one percentage-point increase in the variables included in the VAR. The interest on the sum of coefficients and long-run coefficients follows the literature on TVP-SV-VAR models, starting with Primiceri (2005) and including Canova and Gambetti (2009), among others. These coefficients are viewed as the correct empirical benchmark for detecting violations of the so-called Taylor principle, derived by the theoretical literature. Moreover, two additional reasons justify the interest in long-run responses. The first is an economic one and deals with the fact that central banks might not observe precisely data from the current quarter and thus rather put substantial weight on data from the previous period. The second is an econometric one and it recognizes the fact that in autoregressive models the coefficients on the lags of a process can compensate each other. Concentrating on a single coefficient would thus be misleading if one seeks to explore the contribution of a given variable to changes in the interest rate path.

3 Results

In this section we present estimates for the time-varying coefficients on real house price and real stock price growth as well as for the coefficients on the standard objectives of monetary policy in the context of our baseline model. Later on we perform sensitivity analysis to discuss issues related to simultaneity and to the importance of time-varying parameters and stochastic volatility for our results.

The coefficients are reported along with the 16%-84% credibility intervals for the period ranging from 1985:Q3 to 2008:Q4 since we discard the observations used in the training sample to set up the priors.

3.1 Baseline Estimation

The coefficient on the lagged interest rate finds a counterpart in the interest rate smoothing term in a Taylor-type monetary policy rule. Our median estimate is centered around 0.85, as shown in the top left panel of Figure 1. Incidentally, this value is fully in line with estimates in the DSGE literature. Indeed, the posterior mode estimate for the interest rate parameter is 0.81 in the seminal paper by Smets and Wouters (2007).

In the mid panel in Figure 1 we plot time-varying estimates for the sum of coefficients on current and lagged core PCE inflation and output gap. When compared with Taylor rule coefficients, our estimates appear to be particularly small. The coefficients generally considered standard in a Taylor (1993) rule are higher than 1.5 for the term on inflation and between 0.5 and 1 for the term on the output gap. However, those numbers ignore the role of the interest rate smoothing coefficient. To give some purely illustrative guidance on how to interpret our estimated coefficients, consider as an example the following Taylor-type rule with constant coefficients featuring interest rate smoothing:

$$R_t = (1 - \rho_R)(c + \omega_\pi \pi_t + \omega_y \tilde{y}_t) + \rho_R R_{t-1}, \tag{3.1}$$

where the parenthesis contains the inflation and output gap objectives and c represents a constant, possibly related to the natural rate of interest. As highlighted by the previous equation, our estimated coefficients on inflation and on the output gap should be divided by $(1 - \rho_R)$ to recover average values for ω_π and ω_y of approximately 1.75 and 1.2 respectively.

The coefficients are fairly stable over time, with evidence of a gradual decrease and subsequent increase in the inflation coefficient from early 1992 to late 1998. The output coefficient features the inverse pattern and thus (partially) compensates for the decline in the inflation coefficient. While we find some time variation during the 1990s, the evidence for the 2000s favors more stable coefficients. Nevertheless, we find a smaller response to inflation and higher response to the output gap during the pre-Great Recession period. We will discuss further the implications of our results for the debate over the stance of monetary policy in the pre-Great Recession period in Section 4.

Next, we answer the main questions of interest to this paper, namely i) whether the Fed responded systematically to asset prices and ii) whether this response changed over time. The evidence shown in the bottom panel Figure 1 provides a positive answer to both questions. Indeed, the real S&P 500 growth coefficient is significant and equal to 0.02 around the 1987 financial crisis and otherwise (weakly) insignificant. The magnitude of the coefficient is in the ballpark of the one found by Rigobon and Sack (2003). Moreover, our results are in line with Furlanetto (2011), who shows that the positive and significant coefficient found by Rigobon and Sack (2003) relies on the 1987 financial crisis period being present in the sample. A very similar value (around 0.025, once the estimate for the smoothing coefficient has been taken into account) is found also by Castelnuovo and Nisticò (2010) in an estimated DSGE model with overlapping generations using data on stock price growth as an additional observable variable. All in all, we conclude that stock price growth entered the central bank's reaction function with a statistically significant coefficient only around the 1987 stock market crash.

While our results on the response to stock prices are in line with the previous literature, even though the methodology is different, we uncover some new findings when we investigate the response to real house price growth, which we plot in the bottom panel of Figure 1. The estimated coefficient is significant and roughly equal to 0.1, which is about one third of the inflation coefficient. Such a high response to house prices is estimated over most of our sample with the important exception of the second part of the 1990s where we identify a lower response to house prices (and to PCE inflation). The coefficient gradually decreases from early 2004 and starts rising again back to its previous level in the year prior to the onset of the financial crisis in 2007:Q4. Notably, the decline in the response to house price inflation during the pre-Great Recession period is substantially larger than the decline in the response to consumer price inflation.

We believe that our results on the Fed's conduct are open to different interpretations. On the one hand, we find that the Fed has on average responded to fluctuations in house prices and that this response has on average been quantitatively important. On the other hand, the response declined somewhat sharply precisely in the period when house prices were growing most, i.e. the pre-Great Recession period.

A positive and significant response to house prices from the monetary policy authority to the best of our knowledge has never been found in the VAR literature. As far as we know, only one study provides estimates of the monetary policy response to house prices: Finocchiaro and von Heideken (2013) estimate a DSGE model with nominal loans and collateral constraints in which they embed a monetary policy rule with a direct response to house prices. The mean of their estimated coefficient for the US is 0.36 which, once adjusted for the estimated degree of interest rate smoothing (0.71 in their case), corresponds to the average response over our sample period.

It is important to highlight that our model features a time-varying constant and, thus, a time-varying unconditional mean $\mathbb{E}(\boldsymbol{x}_t) = \left(\mathbf{I} - \mathbf{B}_{1,t}\right)^{-1} \mathbf{c}_t$. The latter may be related to a measure of the natural nominal rate of interest, which is often considered a fixed number in the literature on Taylor rules. In the top right panel of Figure 1 we plot the evolution of the time-varying constant as estimated in our model. It peaks at around 8 percent in 1993 and the declines substantially until the beginning of the new century. Since then it has been reasonably stable, reaching its highest value around 5.75 percent in 2006. As for the other parameters, most time variation seems to be concentrated over the 1990s. Finally, a word of caution regarding these estimates. Fluctuations in the constant in the VAR may reflect fluctuations in the natural rate of interest but also shifts in the inflation target. Our model is unable to disentangle these two sources of variation.

Next, we report in Figure 2 the long-run coefficients which, as in Canova and Gambetti (2009) and Primiceri (2005) among others, measure the total increase in the federal funds rate that would follow a permanent one percentage point increase in the respective variable. Not surprisingly, the dynamics of the long-run coefficients are very similar to the ones presented previously for the sum of coefficients. The only difference lies in the magnitude of the coefficient, which is amplified by the interest rate smoothing term.

While the focus of our paper is the systematic component of monetary policy, an interesting by-product of our analysis is the impulse responses to a monetary policy shock. In order to obtain the impulse responses for the data in terms of levels, we apply the standard practice of cumulating the impulse responses over horizons at each point in time. The magnitude of the impulse responses can thus be interpreted as a percentage change in a given variable in levels following a 1 percentage point increase in the federal funds rate. In Figure 3 we see that a contractionary monetary policy shock has a negative effect on real economic activity as in Gali and Gambetti (2015) and Prieto et al. (2016). Moreover, as in the latter papers and in Primiceri (2005), we do not find a price puzzle in the inflation response over the entire sample period. Turning to the response of stock prices, the latter increase permanently in response to a contractionary monetary policy shock, which is fully in line with the results reported in Gali and Gambetti (2015) for the post-1985 period. Clearly, a positive response of stock prices following a monetary policy shock is at odds with what the "conventional" view would predict, but it can be rationalized in the presence of a bubble component driving stock prices as pointed out theoretically by Galí (2014) and shown empirically by Gali and Gambetti (2015). Finally, a contractionary monetary policy shock unambiguously lowers house prices over the entire sample period, consistent with several previous papers summarized in Williams (2015). Note that the estimated effect is particularly small, at the lower bound of the estimates reviewed in Williams (2015). This is not so surprising, however, since in our model the systematic response of monetary policy largely undoes the direct effect of a monetary policy shock on house prices. We conclude that, according to our model, monetary policy surprises are not effective in curbing house prices.

3.2 Sensitivity Analysis

In this section we perform robustness checks on the estimates for the systematic response of monetary policy to house prices and stock prices.

In the first exercise we change the ordering of the variables in the econometric model. In our baseline we have restricted the impact response of house prices and stock prices to a monetary policy shock since our main focus is on the response of monetary policy to financial variables. Here, we change the vector $\boldsymbol{x}_t$ from $\boldsymbol{x}_t = [\Pi_t \ \widetilde{Y}_t \ \Delta H_t \ \Delta S\&P_t^{500} FFR_t]'$ to $\boldsymbol{x}_t = [\Pi_t \ \widetilde{Y}_t \ FFR_t \ \Delta H_t \ \Delta S\&P_t^{500}]'$, thus imposing that monetary policy does not respond to house prices and stock prices within the quarter. The results, reported in the top panel of Figure 4, point towards a level shift in the coefficients on house prices and S&P 500 but do not change our main result that the FED responded to house prices and that it did so in a time-varying fashion. The response to stock prices is now slightly larger and statistically significant over almost the entire sample period. These results show that different impact responses of variables to shocks (as determined by a different order of the variables) do not have large effects on the estimates for the systematic component of monetary policy. Since the choice of the order is necessarily arbitrary when dealing with fast moving variables such as interest rates, stock prices and (to a lesser extent) house prices, it is reassuring that the main patterns are confirmed.

In a second exercise we use the order $\boldsymbol{x}_t = [\Pi_t \; \widetilde{Y}_t \; \Delta H_t \; FFR_t \Delta S\&P_t^{500}]'$, where we recognize that stock prices are a fast moving variable but maintain the assumption that house prices do not respond within a quarter to monetary policy shocks, as in the baseline. Not surprisingly, the results presented in the mid panel of Figure 4 are an intermediate case between the alternative model presented in the top panel and the baseline model.

In a third exercise we extend the sample to 2015:Q2 by using the Wu and Xia (2016) shadow rate as the monetary policy tool. As is clearly evident from the results reported in the bottom panel of Figure 4, we can not overturn the conclusion that the coefficients on house prices and stock prices were positive and significant prior to the onset of the Great Recession and around the 1987 stock market crash, respectively.

The last three exercises are devoted to understanding the role of time variation and stochastic volatility in our econometric model. As becomes evident from Figure 5, it is crucial to account for both channels. Indeed, shutting down one or both of them would lead a researcher to mistakenly infer that the FED did not care about house prices in its conduct of monetary policy. While the response to stock prices is preserved in the absence of stochastic volatility, this is not the case for house prices. To detect a positive response to house prices, both time variation and stochastic volatility are needed. This result justifies the use of a model with time variation and stochastic volatility to analyze our research question, rather than a simple constant coefficient VAR, even if the estimated amount of time variation is rather limited (although non-negligible). Note that a limited amount of time variation in the estimated coefficients is in keeping with the previous literature using this kind of model (cf. Belongia and Ireland (2016b), Canova and Gambetti (2009) and Primiceri (2005)).

4 Debate on the Evolution of US Monetary Policy

In this section we discuss the evolution of US monetary policy in the pre-Great Recession period. In particular, we contribute to the debate between Taylor (2007, 2009) and Bernanke (2015) through the lenses of our model that includes house prices and stock prices.

In recent years, Taylor argued that the FOMC policy has been "too low for too long" compared to the interest rate path prescribed by his rule (Taylor (2007, 2009)). Bernanke (2015) shows that when using i) real-time data, ii) a modified Taylor rule with a coefficient of 1 for the output gap and iii) PCE inflation instead of GDP inflation, the "Great Deviation" pointed out by Taylor does not emerge.

Belongia and Ireland (2016b) estimate a three variable TVP-SV-VAR model and find evidence of i) declining coefficients in their model's estimated policy rule, pointing to a shift in the Fed's emphasis away from stabilizing inflation during the period 2000-2007, and ii) large expansionary monetary policy shocks during the period 2003:Q3-2004:Q2, seen as evidence of a more discretionary policy and thus supporting Taylor's argument.

Our model with house prices and stock prices confirms some limited changes in the systematic component of monetary policy. We find evidence of a slightly lower response to inflation (but not to the output gap) in the pre-Great Recession period. Moreover, we also find a decline in the response to house prices that points to a less aggressive reaction of monetary policy to economic conditions. However, we find only relatively small changes in the non-systematic component of monetary policy. More specifically, we consider the fitted interest rate implied by our model:

$$\begin{aligned}\hat{R}_t = & - \hat{a}_{51,t}\Pi_t - \hat{a}_{52,t}\widetilde{Y}_t - \hat{a}_{53,t}\Delta H_t - \hat{a}_{54,t}\Delta S\&P^{500}_t \\ & + \widehat{ab}^1_{51,t}\Pi_{t-1} + \widehat{ab}^1_{52,t}\widetilde{Y}_{t-1} + \widehat{ab}^1_{53,t}\Delta H_{t-1} + \widehat{ab}^1_{54,t}\Delta S\&P^{500}_{t-1} + \widehat{ab}^1_{55,t}R_{t-1},\end{aligned}$$

The top panel of Figure 7 shows that the implied VAR interest rate fits well with the actual interest rate movements. We find some expansionary monetary policy shocks during the period 2003-2004, but we do not detect a "Great Deviation" from the historical rule. In Figure 6 we plot the posterior median of the standard deviation of monetary policy shocks and we note a period of relatively high volatility at the beginning of the new century. However, this volatility declines rapidly from 2003 and reaches its minimum in 2004 before increasing again. Moreover, the expansionary shocks estimated over the period 2003-2004 are somewhat balanced by a series of contractionary shocks in 2006-2007.

Overall, we confirm previous results in Belongia and Ireland (2016b), as we identify a shift in Federal Reserve policy away from inflation and house price inflation stabilization and some departures from rule-like behavior. However, our reading of the results is that these changes are quantitatively small and that this evidence is not sufficient to conclude that the interest rate was too low for too long.

Next, Orphanides (2001) stresses that real-time policy recommendations may differ considerably from those obtained with ex-post revised data. However, Croushore and Evans (2006) show that accounting for data revisions has only a modest effect quantitatively on the recursively identified monetary policy shock measures and impulse responses obtained from standard VARs, supporting the common approach of using ex-post revised data in structural VARs for monetary policy analysis. Nevertheless, to ensure that our estimated time-varying parameters are also relevant in real time, we plot in the bottom panel of Figure 7 the implied interest rate path based on the estimated TVP-SV-VAR parameters, replacing the ex-post values for core PCE inflation and the output gap with their real-time counterparts. We obtained the real-time data vintages from the Federal Reserve Bank of Philadelphia's Real-Time Dataset for Macroeconomists (RTDSM), described in Croushore and Stark (2001).[4] The bottom panel of Figure 7 shows that also when using real-time data the implied VAR interest rate fits well with the actual interest rate movements. In keeping with the argument in Bernanke (2015), the use of real-time data reduces further the role of unsystematic policy in recent years.

[4]Note that real-time data vintages for PCE inflation are only available from 1996:Q2 onwards.

To sum up, in the case of both real-time data and ex-post data, we do not find evidence of a "Great Deviation" from the path prescribed by the systematic component estimated by our model, in line with Bernanke's arguments.

5 Conclusion

The main contribution of this paper is to provide evidence on the time-varying response of monetary policy to stock prices and house prices. We find that the response to stock price fluctuations has been small and episodic, in keeping with the previous literature. Our main result is that we find a significant response to house prices, both in economic and statistical terms. While the response to house prices declines somewhat in the pre-Great Recession period, our evidence shows that the Fed considers variables other than inflation and real economic activity in its estimated reaction function. Our analysis has no normative implications for whether such a response to asset prices (and house prices in particular) was optimal, insufficient or excessive. Nevertheless, we believe it is interesting to document that it was substantial.

One direction for future research is to further take into account the simultaneity in the determination of house prices, stock prices and interest rates. In this paper we have shown that our results hold when imposing different orders. However, it would be interesting to explore alternative identification schemes, perhaps building on previous attempts to deal with the simultaneity problem in the context of models with constant coefficients (cf. Bjørnland and Leitemo (2009), D'Amico and Farka (2011), Rigobon and Sack (2003)). Bringing these insights into models with TVP-SV-VAR seems an interesting avenue for future research.

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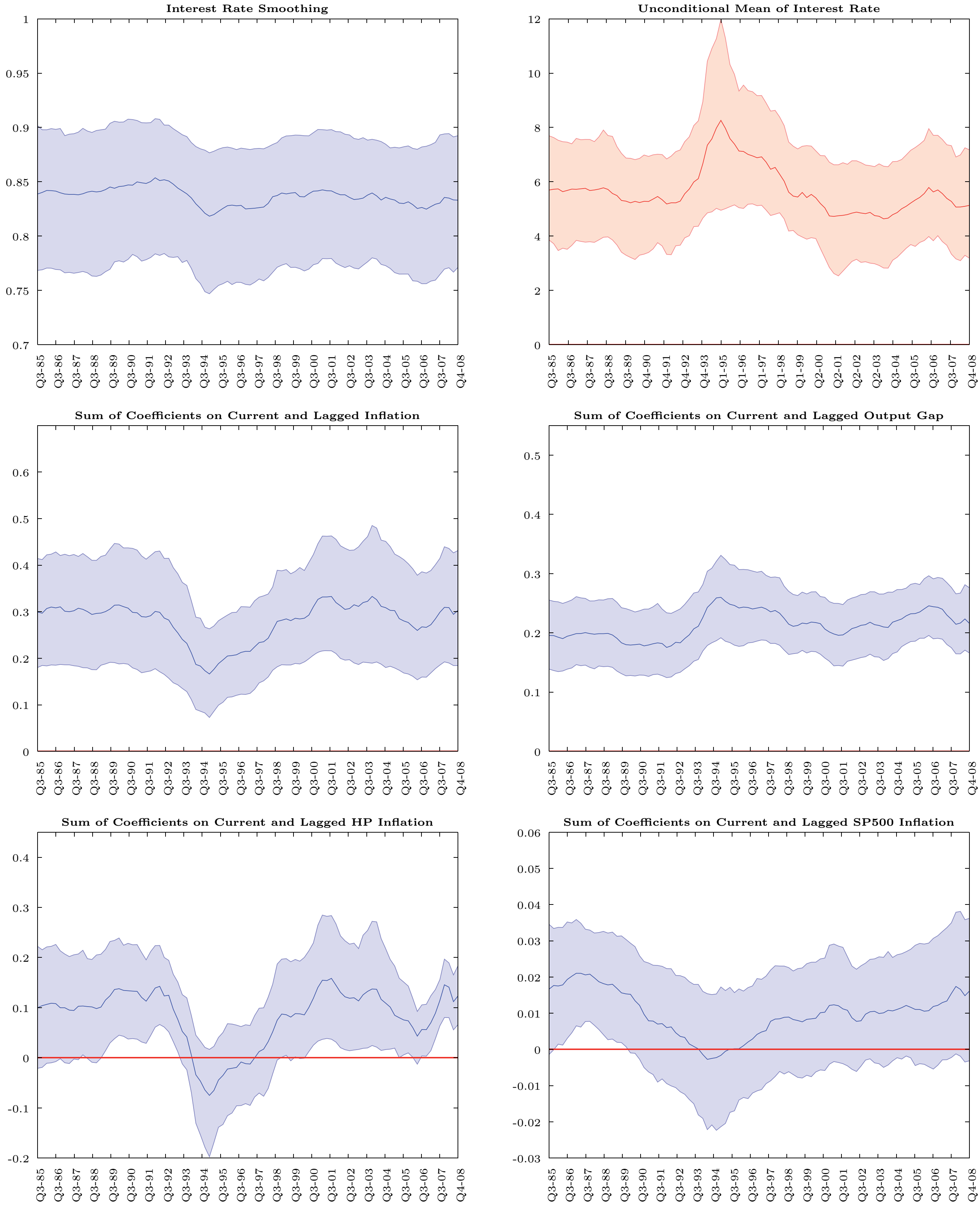

Figure 1: *Top Panel*: Degree of Interest Rate Smoothing and Interest Rate Trend, *Mid Panel*: Sum of Coefficients on Inflation and Output Gap, *Bottom Panel*: Sum of Coefficients on Real House Price Inflation and S&P 500 Growth.

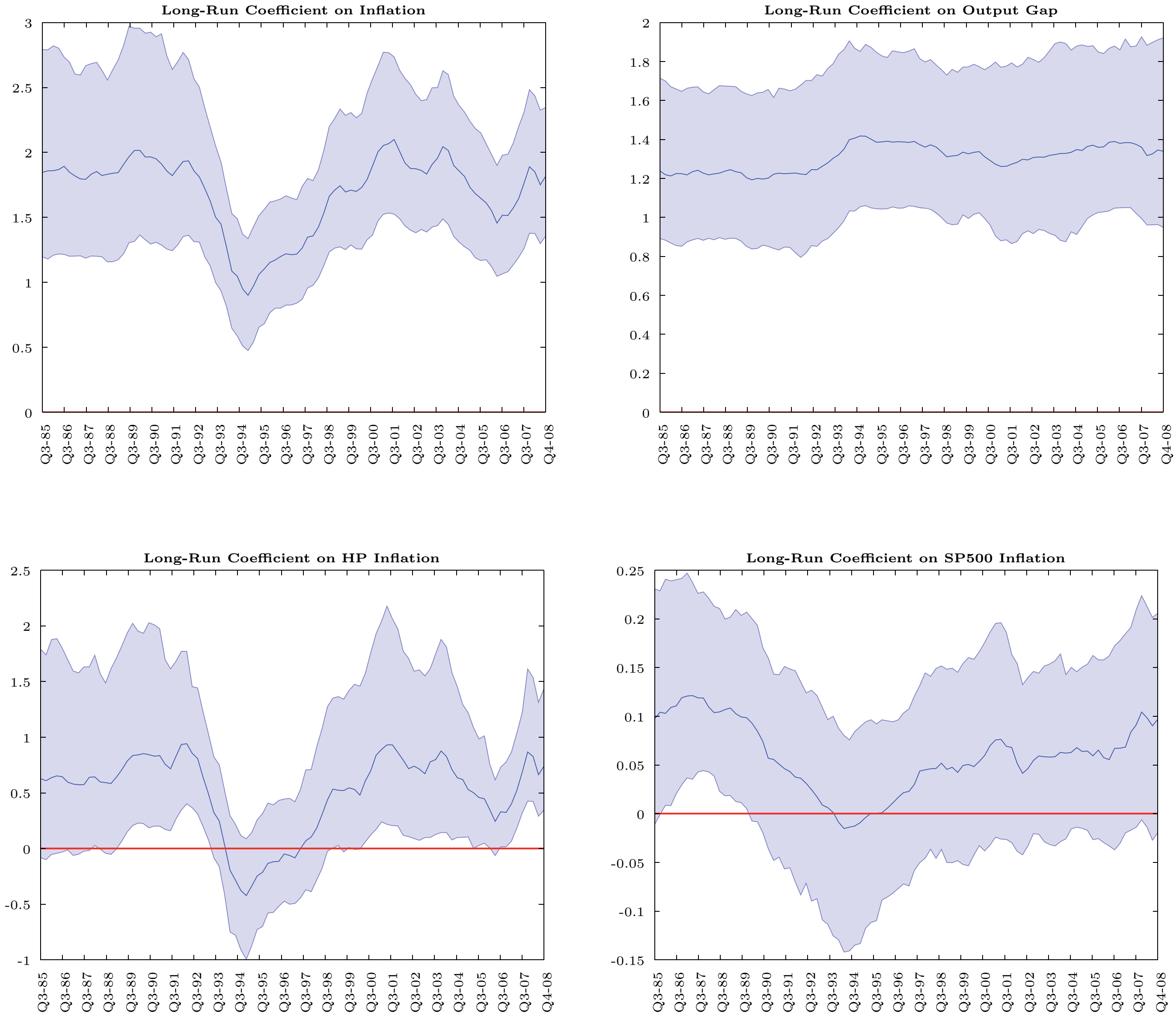

Figure 2: Long-Run Coefficients. *Top Panel*: Inflation and Output Gap, *Bottom Panel*: Real House Price Inflation and S&P 500 Growth.

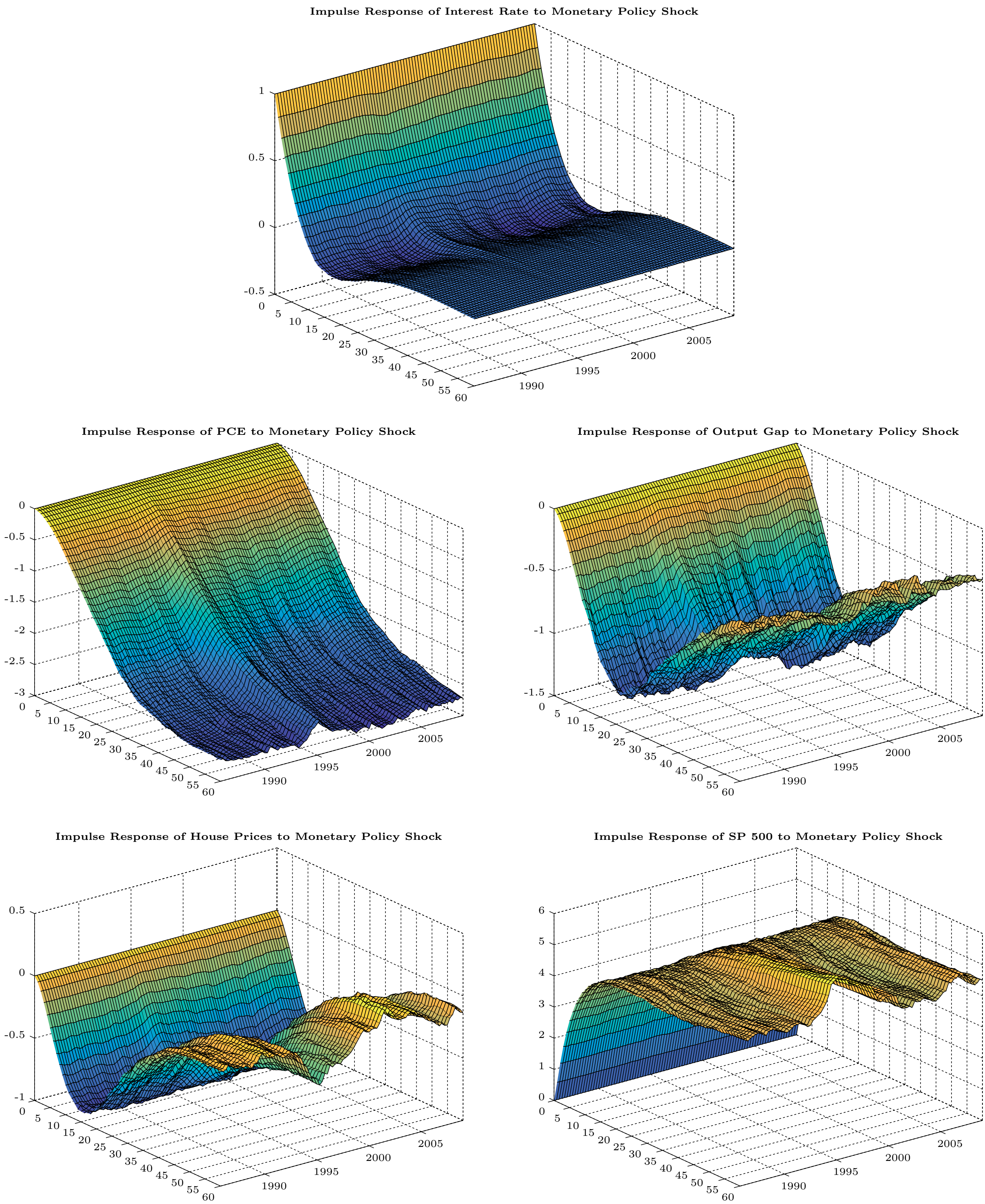

Figure 3: Impulse Responses to a Monetary Policy Shock. *Top Panel*: Interest Rate, *Mid Panel*: PCE Prices and Output Gap, *Bottom Panel*: House Prices and Stock Prices.

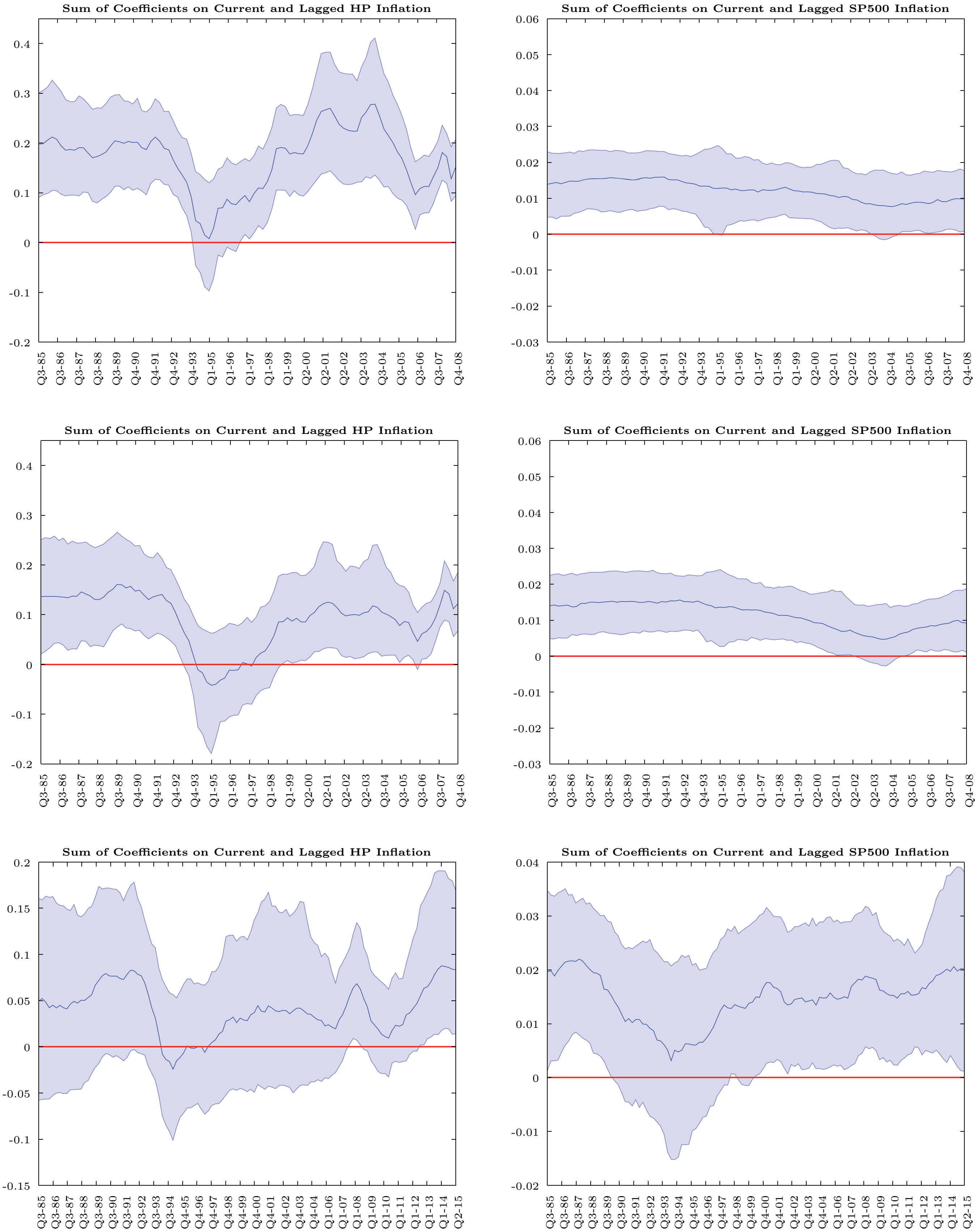

Figure 4: Sum of Coefficients on Real House Price Inflation and S&P 500 Growth. *Top Panel*: Reordered Model with $\boldsymbol{x}_t = [\Pi_t \ \widetilde{Y}_t \ FFR_t \ \Delta H_t \ \Delta S\&P_t^{500}]'$ (Top Panel), *Mid Panel*: Reordered Model with $\boldsymbol{x}_t = [\Pi_t \ \widetilde{Y}_t \ \Delta H_t \ FFR_t \ \Delta S\&P_t^{500}]'$, *Bottom Panel*: Model Extended to 2015:Q2, Shadow Rate as Monetary Instrument.

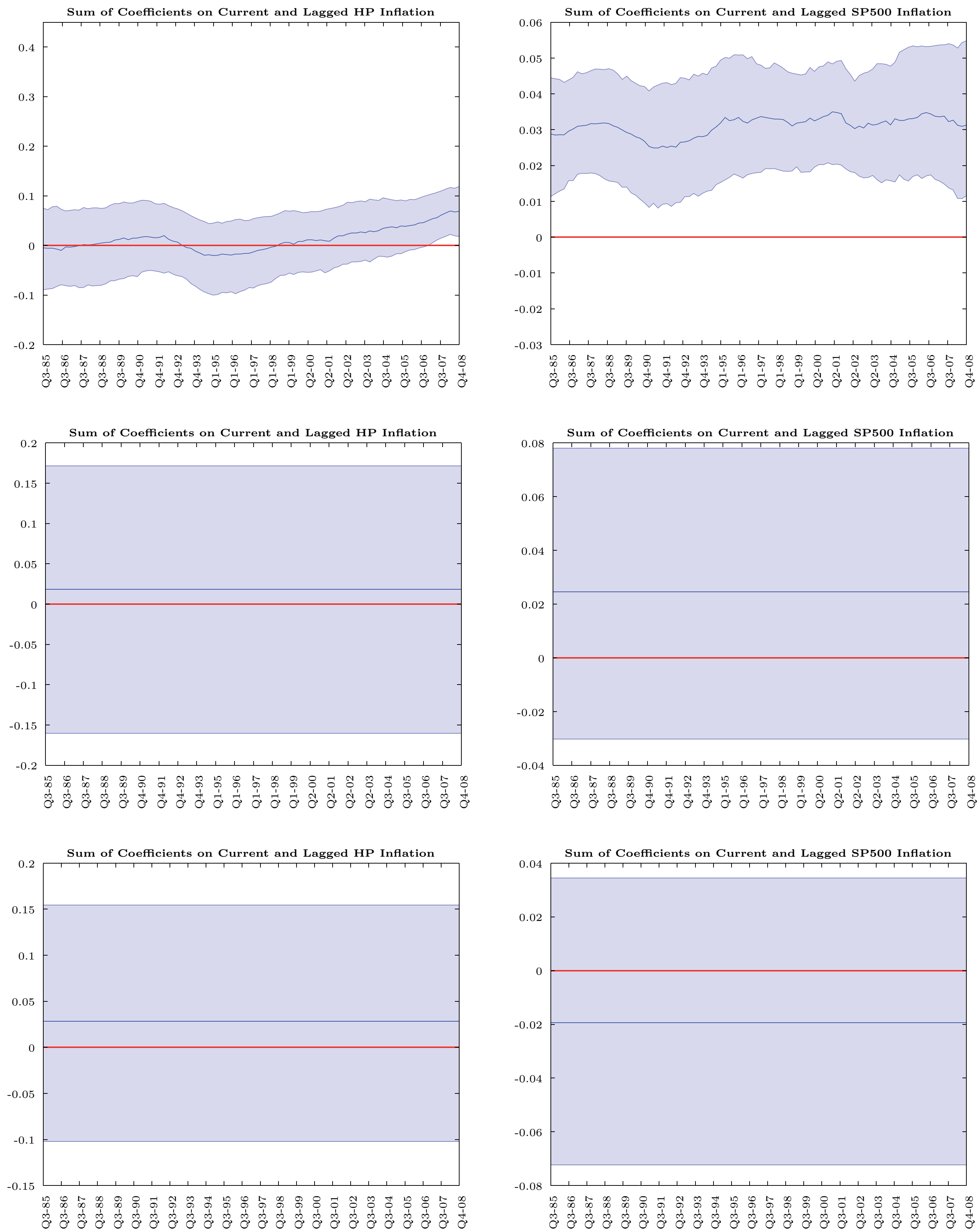

Figure 5: Sum of Coefficients on Real House Price Inflation and S&P 500 Growth. *Top Panel*: Model with No Stochastic Volatility, *Mid Panel*: Model with No Time Variation in A_t and $B_{1,t}$, *Bottom Panel*: Model with No Stochastic Volatility and No Time Variation in A_t and $B_{1,t}$.

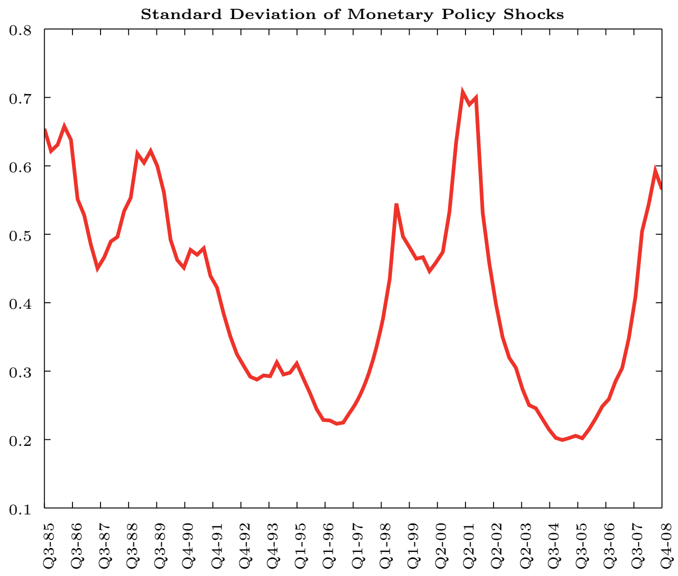

Figure 6: Standard Deviation of Monetary Policy Shocks Over Time.

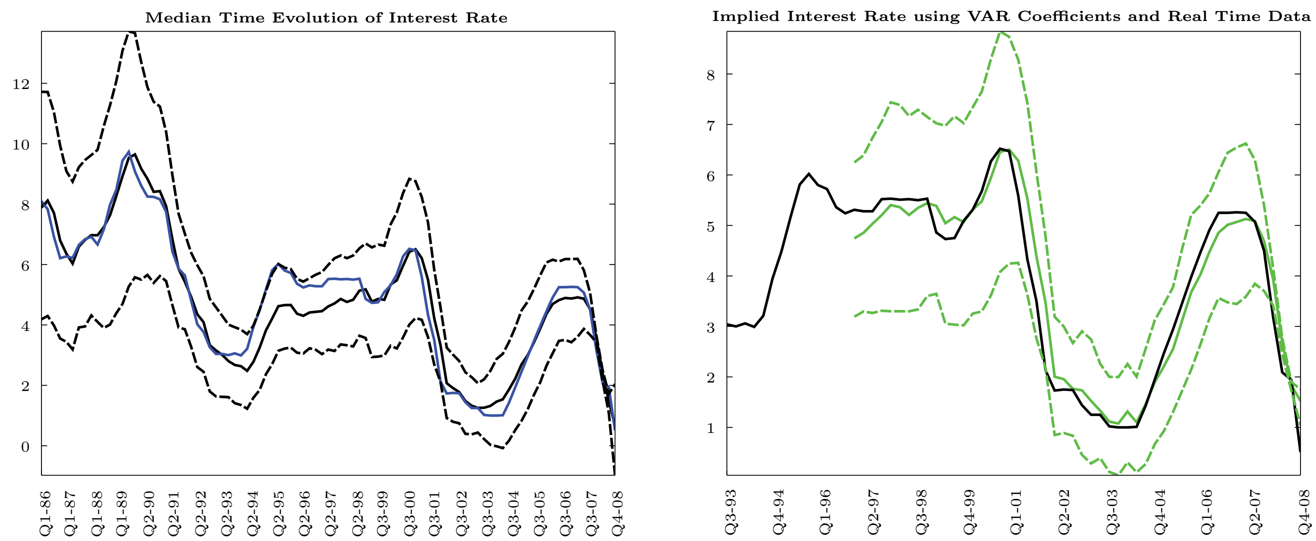

Figure 7: *Left Panel*: VAR Implied Interest Rate (black) vs Actual Interest Rate (blue). *Right Panel*: VAR Implied (green) vs Actual Interest Rate (black), Real-Time Data.

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