The Effects of Monetary Policy on Asset Prices Bubbles: Some Evidence *

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Abstract

We estimate the response of stock prices to exogenous monetary policy shocks using vector-autoregressive models with time-varying parameters. Under our baseline identification scheme, the evidence cannot be easily reconciled with conventional views on the effects of interest rate changes on asset price bubbles.

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The economic and financial crisis of 2008-2009 has been associated in many countries with a rapid decline in housing prices, following a protracted real estate boom. This has attracted a renewed interest in the link between monetary policy and asset price bubbles, and revived the long standing debate on whether and how monetary policy should respond to perceived deviations of asset prices from fundamentals.[1]

The consensus view before the crisis was that central banks should focus on stabilizing inflation and the output gap, and ignore fluctuations in asset prices or potential bubbles.[2] The recent crisis has challenged that consensus and has strengthened the viewpoint that central banks should pay attention and eventually respond to developments in asset markets. Supporters of this view argue that monetary authorities should "lean against the wind," i.e. raise the interest rate to counteract any bubble-driven episode of asset price inflation, even at the cost of temporarily deviating from their inflation or output gap targets. Any losses associated with these deviations, it is argued, would be more than offset by the avoidance of the consequences of a future burst of the bubble.[3]

A central tenet of the case for "leaning against the wind" monetary policies is the presumption that an increase in interest rates will reduce the size of an asset price bubble. While that presumption may have become part of the received wisdom, no empirical or theoretical support seems to have been provided by its advocates.

In recent work (Galí (2013)), one of us has challenged, on theoretical grounds, the link between interest rates and asset price bubbles maintained by the conventional view. The reason is that, at least in the case of *rational* asset price bubbles, the bubble component must grow, in equilibrium, at the rate of interest. If that is the case, an interest rate increase may end up enhancing the size of the bubble. Furthermore, and as discussed below, the theory of rational bubbles implies that the effects of monetary policy on *observed* asset prices depend on the relative size of the bubble component. More specifically, an increase in the interest rate should have a negative impact on the price of an asset in periods where the bubble component is small compared to the fundamental. The reason is that an interest rate increase always reduces the "fundamental" price of the asset, an effect that should be dominant in "normal" times, when the bubble component is small or nonexistent. But if the relative size of the bubble is large, an interest rate hike may end up increasing the

[1]Throughout the paper we use the term "monetary policy" in the narrow sense of "interest rate policy." Thus we exclude from that definition policies involving macroprudential instruments which are sometimes controlled by central banks and which may also be used to stabilize asset prices.

[2]See, e.g. Bernanke and Gertler (1999, 2000) and Kohn (2006). Two arguments have been often pointed to in support of that view: (i) asset price bubbles are difficult to detect and measure, and (ii) interest rates are "too blunt" an instrument to prick a bubble, and their use with that purpose may have unintended collateral damages.

[3]See, e.g., Borio and Lowe (2002) and Cecchetti et al. (2000) for an early defense of "leaning against the wind" policies.

observed asset price over time, due to its positive effect on the bubble more than offsetting the negative impact on the fundamental.

In the present paper we provide evidence on the response of stock prices to monetary policy shocks, and try to use that evidence to infer the nature of the impact of interest rate changes on the (possible) bubble component of stock prices..Our main goal is to assess the empirical merits of the "conventional" view. Under the latter, the size of the bubble component of stock prices should decline in response to an exogenous increase in interest rates. Since the fundamental component is expected to go down in response to the same policy intervention, any evidence pointing to a positive response of *observed* stock prices (i.e. of the sum of the fundamental and bubble components) to an exogenous interest rate hike would call into question the conventional view regarding the effects of monetary policy on stock price bubbles.

Our starting point is an estimated vector-autoregression (VAR) on quarterly US data for GDP, inflationm, dividends, the federal funds rate, and a stock price index (S&P500). Our identification of monetary policy shocks is based on the approach of Christiano Eichenbaum and Evans (2005; henceforth, CEE), though our focus here is on the dynamic response of stock prices to an exogenous hike in the interest rate. Also, and in contrast with CEE, we allow for time-variation in the VAR coefficients, which results in estimates of time-varying impulse responses of stock prices to policy shocks.[4] In addition to the usual motivations for doing this (e.g. , structural change), we point to a new one which is specific to the issue at hand: to the extent that changes in interest rates have a different impact on the fundamental and bubble components, the overall effect on the observed stock price may change over time as the relative size of the bubble changes.

Under our baseline specification, which assumes no contemporaneous response of monetary policy to asset prices, the evidence points to protracted episodes in which stock prices increase persistently in response to an exogenous tightening of monetary policy. That response is clearly at odds with the "conventional" view on the effects of monetary policy on bubbles, as well as with the predictions of bubbleless models. We also argue that it is unlikely that such evidence be accounted for by an endogenous response of the equity premium to the monetary policy shocks.

When we allow for an endogenous contemporaneous response of interest rates to stock prices, and calibrate the relevant coefficient in the monetary policy rule according to the findings in Rigobon and Sack (2003), our findings change dramatically: stock prices decline substantially in response to a tightening of monetary policy, more so than our estimated fundamental components. That finding would seem to vindicate the conventional view on

[4]See, e.g. Primiceri (2005) and Gali and Gambetti (2009) for some macro applications of the TVC-VAR methodology.

the effectiveness of leaning against the wind policies. Recent evidence by Furlanetto (2011), however, calls into question the relevance of this alternative specification for much of the sample period analyzed, while supporting instead our baseline specification.

Ultimately, our objective is to produce evidence that can improve our understanding of the impact of monetary policy on asset prices and asset price bubbles. That understanding is a necessary condition before one starts thinking about how monetary policy *should* respond to asset prices.

The remainder of the paper is organized as follows. In Section 1 we discuss alternative hypothesis on the link between interest rates and asset prices. Section 2 describes our empirical model. In Section 3 we report the main findings under our baseline specification. Section 4 provides alternative interpretations as well as evidence based on an alternative specification. Section 5 concludes.

1 Monetary policy and asset price bubbles: Theoretical issues

We use a simple partial equilibrium asset pricing model to introduce some key concepts and notation used extensively below.[5] We assume an economy with risk neutral investors and an exogenous, time-varying (gross) riskless interest rate R_t.[6] Let Q_t denote the price in period t of an infinite-lived asset, yielding a dividend stream $\{D_t\}$.

We interpret that price as the sum of two components: a "fundamental" component, Q_t^F, and a "bubble" component, Q_t^B. Formally,

$$Q_t = Q_t^F + Q_t^B \tag{1}$$

where the *fundamental* component is defined as the present discounted value of future dividends:

$$Q_t^F \equiv E_t \left\{ \sum_{k=1}^{\infty} \left(\prod_{j=0}^{k-1} (1/R_{t+j}) \right) D_{t+k} \right\} \tag{2}$$

or, rewriting it in log-linear form (and using lower case letters to denote the logs of the original variables)[7]

$$q_t^F = const + \sum_{k=0}^{\infty} \Lambda^k \left[(1-\Lambda) E_t\{d_{t+k+1}\} - E_t\{r_{t+k}\} \right] \tag{3}$$

where $\Lambda \equiv \Gamma/R < 1$, with Γ and R are denoting, respectively, the (gross) rates of dividend growth and interest along a balanced growth path.

[5] See Galí (2013) for a related analysis in general equilibrium.

[6] Below we discuss the implications of allowing for a risk premium.

[7] See, e.g., Cochrane (2001, p.395) for a derivation.

How does a change in interest rates affect the price of an asset that contains a bubble? We can seek an answer to that question by combining the dynamic responses of the two components of the asset price to an exogenous shock in the policy rate. Letting that shock be denoted by ε_t^m, we have:

$$\frac{\partial q_{t+k}}{\partial \varepsilon_t^m} = (1-\gamma_{t-1})\frac{\partial q_{t+k}^F}{\partial \varepsilon_t^m} + \gamma_{t-1}\frac{\partial q_{t+k}^B}{\partial \varepsilon_t^m} \tag{4}$$

where $\gamma_t \equiv Q_t^B/Q_t$ denotes the share of the bubble in the observed price in period t.

Using (2), we can derive the predicted response of the fundamental component

$$\frac{\partial q_{t+k}^F}{\partial \varepsilon_t^m} = \sum_{j=0}^{\infty} \Lambda^j \left((1-\Lambda)\frac{\partial d_{t+k+j+1}}{\partial \varepsilon_t^m} - \frac{\partial r_{t+k+j}}{\partial \varepsilon_t^m} \right) \tag{5}$$

Both conventional wisdom and economic theory (as well as the empirical evidence discussed below) point to a rise in the real interest rate and a decline in dividends in response to an exogenous tightening of monetary policy, i.e. $\partial r_{t+k}/\partial \varepsilon_t^m > 0$ and $\partial d_{t+k}/\partial \varepsilon_t^m \leq 0$ for $k = 0, 1, 2, ...$Accordingly, the fundamental component of asset prices is expected to decline in response to such a shock, i.e. we expect $\partial q_{t+k}^F/\partial \varepsilon_t^m < 0$ for $k = 0, 1, 2, ...$

Under the "conventional view" on the effects of monetary policy on asset price bubbles we have, in addition:

$$\frac{\partial q_{t+k}^B}{\partial \varepsilon_t^m} \leq 0 \tag{6}$$

for $k = 0, 1, 2, ...$i.e. a tightening of monetary policy should lead, sooner or later, to a decline in the size of the bubble. Hence, the overall effect on the observed asset price should be unambiguously negative, independently of the relative size of the bubble:

$$\frac{\partial q_{t+k}}{\partial \varepsilon_t^m} < 0$$

for $k = 0, 1, 2, ...$

As argued in Galí (2013), however, the premise of a decline in the size of the bubble in response to a interest rate hike does not have a clear theoretical underpinning. In fact, the theory of *rational* asset price bubbles opens the door for a very different prediction. To see this, note that the following difference equation must hold in a rational expectations equilibrium:

$$Q_t R_t = E_t\{D_{t+1} + Q_{t+1}\} \tag{7}$$

It can be easily checked that (2) satisfies

$$Q_t^F R_t = E_t\{D_{t+1} + Q_{t+1}^F\} \tag{8}$$

Using (1), (7) and (8), it can be easily checked that the bubble component must satisfy:[8]

$$Q_t^B R_t = E_t\{Q_{t+1}^B\} \tag{9}$$

or, equivalently, in its log-linear version:

$$E_t\{\Delta q_{t+1}^B\} = r_t$$

Hence, an increase in the interest rate will raise the expected growth of the bubble component. Note that the latter corresponds to the bubble's expected return, which–under the risk neutrality assumption made here–must be equal to the interest rate. Accordingly, and as discussed in Galí (2013), any rule that implies a systematic positive response of the interest rate to the size of the bubble, will tend to amplify the movements in the latter –an outcome that calls into question the conventional wisdom about the relation between interest rates and bubbles.

Changes in interest rates, however, may also affect the bubble through a second channel: a possible systematic comovement between the (*indeterminate*) innovation in the bubble with the surprise component of the interest rate. To see this, evaluate the previous expression at $t-1$ and eliminate the expectational operator to obtain:

$$\Delta q_t^B = r_{t-1} + \xi_t \tag{10}$$

where $\xi_t \equiv q_t^B - E_{t-1}\{q_t^B\}$ is an arbitrary process satisfying $E_{t-1}\{\xi_t\} = 0$ for all t (i.e. the martingale-difference property). Note that the unanticipated change ("innovation") in the size of the bubble, ξ_t, may or may not be related to fundamentals and, in particular, to the interest rate innovation, $r_t - E_{t-1}\{r_t\}$. Thus, one can write:

$$\xi_t = \psi_t(r_t - E_{t-1}\{r_t\}) + \xi_t^* \tag{11}$$

where ψ_t is a (possibly random) parameter and $\{\xi_t^*\}$ is a zero-mean martingale-difference process orthogonal to interest rate innovations at all leads and lags, i.e. $E\{\xi_t^* r_{t-k}\} = 0$, for $k = 0, \pm 1, \pm 2, ...$ Note that neither the sign nor the size of ψ_t, nor its possible dependence on the policy regime, are pinned down by the theory. Accordingly, the *contemporaneous impact* of an interest rate innovation (or of any other shock) on the bubble is, in principle, indeterminate.

Hence, the dynamic response of the bubble component to a monetary policy tightening is given by

$$\frac{\partial q_{t+k}^B}{\partial \varepsilon_t^m} = \begin{cases} \psi_t \frac{\partial r_t}{\partial \varepsilon_t^m} & \text{for } k = 0 \\ \psi_t \frac{\partial r_t}{\partial \varepsilon_t^m} + \sum_{j=0}^{k-1} \frac{\partial r_{t+j}}{\partial \varepsilon_t^m} & \text{for } k = 1, 2, ... \end{cases} \tag{12}$$

[8] Transversality conditions generally implied by optimizing behavior of infinite-lived agents are often used to rule out such a bubble component (see, e.g., Santos and Woodford (1997)). That constraint does not apply to economies with overlapping generations of finitely-lived agents (e.g., Samuelson (1958), Tirole (1985)).

for $k = 0, 1, 2, ...$Thus, and as discussed above, the initial impact on the bubble, captured by coefficient ψ_t, is indeterminate, both in sign and size. Yet, and conditional on $\partial r_{t+k}/\partial\varepsilon_t^m > 0$, for $k = 0, 1, 2, ..$ the subsequent growth of the bubble is predicted to be positive. The long run impact of the monetary policy shock on the size of the bubble, $\lim_{k\to\infty} \partial q_{t+k}^B/\partial\varepsilon_t^m$ will be positive or negative depending on whether the persistence of the real interest rate response is more than sufficient to offset any eventual negative initial impact. Thus, when considered in combination with the predicted response of the fundamental component, the theory of rational bubbles implies that the sign of the response of observed asset prices to a tightening of monetary policy is ambiguous. Most importantly, however, from the viewpoint of assessing the empirical validity of this perspective, it opens the door to the possibility that the observed asset price rises (possibly after some initial decline), as long as one or more of the following conditions are satisfied: (i) ψ_t is not "too negative", (ii) the response of the real interest rate is persistent enough, and (iii) the relative size of the bubble γ_t is large enough (so that the eventual positive response of the latter more than offsets the likely decline in the fundamental component).

To illustrate the previous discussion, consider an asset whose dividends are exogenous and independent of monetary policy. In response to an exogenous policy tightening the real interest rate is assumed to evolve according to $\partial r_{t+k}/\partial\varepsilon_t^m = \rho_r^k$, for $k = 0, 1, 2, ...$The response of the (log) asset price to a unit shock is then given by:

$$\frac{\partial q_{t+k}}{\partial\varepsilon_t^m} = -(1-\gamma_{t-1})\frac{\rho_r^k}{1-\Lambda\rho_r} + \gamma_{t-1}\left(\psi_t + \frac{1-\rho_r^k}{1-\rho_r}\right)$$

Figure 1 displays the dynamic responses of the asset price for alternative configurations of γ and ψ. In all cases we assume $\Lambda = 0.99$ and $\rho_r = 0.8$. The black line (with circles) displays the asset price response in the absence of a bubble ($\gamma_{t-1} = 0$). The asset price declines on impact, and gradually returns to its original value. The blue line (with circles) shows the response for $\gamma_t = 0.5$ and $\psi_t = 0$. Note that the asset price also declines on impact, but now it recovers at a faster pace (due to the growing bubble) and ends up overshooting permanently its initial value and stabilizing at a higher level. The red line (with squares) corresponds to the case of $\gamma_t = 0.5$ and $\psi_t = -8$. Now the negative impact of the interest rate hike on the asset price is larger, due to its initial shrinking effect on the size of the bubble. Finally, the green line (with triangles) shows the response under $\gamma_t = 0.5$ and $\psi_t = 6$. Now the asset price already rises from the time of the shock, given that the positive response of the bubble on impact more than offsets the decline of the fundamental component.

The previous simulations make clear that the theory of rational bubbles is consistent with a broad range of responses of asset prices to a tightening of monetary policy. By contrast, the conventional view predicts an unambiguous decline in asset prices, for both the fundamental and bubble component are expected to go down in response to a policy tightening. Accordingly, any evidence of a decline in asset prices in response to that tightening would not be conclusive as to the validity of the two views on the effects of monetary policy on the bubble. On the other hand, any evidence of a positive impact on the asset price at some horizon subsequent to the same policy intervention would be clearly at odds with both the key premise and the implications of the "conventional view," while consistent (at least, qualitatively) with the theory of rational bubbles.

2 The empirical model

The present section describes our empirical model, which consists of a structural vector autoregression model with time-varying coefficients (TVC-SVAR). Though focusing on different variables, the specification of the reduced form time-varying VAR follows closely that in Primiceri (2005). On the other hand our choice of variables and identification strategy follows that in Christiano et al. (2005). Our constant coefficients VAR, for which we also report results below, can be seen as a limiting case of the model with time-varying coefficients, so we do not provide a separate description.

Let y_t, p_t, i_t, q_t, and d_t denote, respectively, (log) output, the (log) price level, the short-term nominal interest rate controlled by the central bank, the (log) stock price index, and its corresponding (log) dividend series (both in real terms). We define $x_t \equiv [\Delta y_t, \Delta d_t, \Delta p_t, i_t, \Delta q_t]'$. The relationship between those variables and the structural shocks is assumed to take the form of an autoregressive model with time-varying coefficients:

$$x_t = A_{0,t} + A_{1,t}x_{t-1} + A_{2,t}x_{t-2} + ... + A_{p,t}x_{t-p} + u_t \tag{13}$$

where $A_{0,t}$ is a vector of time-varying intercepts, and $A_{i,t}$, for $i = 1, ...,$and S_t are matrices of time-varying coefficients, and where the vector of reduced form innovations u_t follows a white noise Gaussian process with mean zero and covariance matrix Σ_t. We assume the reduced form innovations are a linear transformation of the underlying structural shocks ε_t given by

$$u_t \equiv S_t \varepsilon_t$$

where $E\{\varepsilon_t \varepsilon_t'\} = I$ and $E\{\varepsilon_t \varepsilon_{t-k}'\} = 0$ for all t and $k = 1, 2, 3, ...$

Let $\theta_t = vec(A_t')$ where $A_t = [A_{0,t}, A_{1,t}..., A_{p,t}]$ and $vec(\cdot)$ is the column stacking operator. We assume θ_t evolves over time according to the process

$$\theta_t = \theta_{t-1} + \omega_t \tag{14}$$

where ω_t is a Gaussian white noise process with zero mean and constant covariance Ω, and independent of u_t at all leads and lags.

We model the time variation of Σ_t as follows. Let $\Sigma_t = F_t D_t F_t'$, where F_t is lower triangular, with ones on the main diagonal, and D_t a diagonal matrix. Let σ_t be the vector containing the diagonal elements of $D_t^{1/2}$ and $\phi_{i,t}$ a column vector with the non-zero elements of $(i+1)th$ row of F_t^{-1}. We assume that

$$\log \sigma_t = \log \sigma_{t-1} + \xi_t \tag{15}$$

$$\phi_{i,t} = \phi_{i,t-1} + \upsilon_{i,t} \tag{16}$$

where ξ_t and $\upsilon_{i,t}$ are white noise Gaussian processes with zero mean and (constant) covariance matrices Ξ and Ψ_i, respectively. We assume that $\upsilon_{i,t}$ is independent of $\upsilon_{j,t}$, for $j \neq i$, and that ω_t, ε_t, ξ_t and $\upsilon_{i,t}$ (for $i = 1, ..., n-1$) are mutually uncorrelated at all leads and lags. Note that the constant coefficient VAR can be seen as a limiting case of the previous model with $\Omega = 0, \Xi = 0, \Psi_i = 0$.

Our identification of the monetary policy shock is inspired by the strategy proposed by Christiano, Eichenbaum and Evans (2005). More specifically we assume that the monetary policy shock does not affect GDP, dividends or inflation contemporaneously. In addition, our baseline specification assumes that the central bank does not respond contemporaneously to innovations in real stock prices.[9] Letting the fourth element in ε_t, denoted by ε_t^m, correspond to the monetary policy shock, the first assumption implies that the fourth column of S_t has zeros as its first three elements, while its two remaining elements are unrestricted. The second assumption implies that the last element in the fourth row of S_t is zero. Since our focus is on monetary policy shocks, we need not place any other restrictions on matrix S. To facilitate implementation we just let S_t be the Cholesky factor of Σ_t, i.e. the unique lower triangular matrix such that $S_t S_t' = \Sigma_t$, but make no attempt to interpret the remaining "structural" shocks. [to be added: robustness to alternative values of $[S_t]_{4,5}$].

To define the impulse response functions let us rewrite (13) in companion form:

$$\mathbf{x}_t = \boldsymbol{\mu}_t + \mathbf{A}_t \mathbf{x}_{t-1} + \mathbf{u}_t$$

where $\mathbf{x}_t \equiv [x_t', x_{t-1}', ..., x_{t-p+1}']'$, $\mathbf{u}_t \equiv [u_t', 0, ..., 0]'$, $\boldsymbol{\mu}_t \equiv [A_{0,t}', 0, ..., 0]'$ and $\mathbf{A}_t$ is the corresponding companion matrix. We use a *local approximation* of the implied dynamic response to a t period shock. Formally, the local response is given by

$$\frac{\partial x_{t+k}}{\partial u_t'} = [\mathbf{A}_t^k]_{5,5} \equiv B_{t,k}$$

for $k = 1, 2, ..$where $[M]_{5,5}$ represents the first 5 rows and 5 columns of any matrix M, and where $B_{t,0} \equiv I$. Thus, the dynamic responses of the variables in x_t to a monetary policy

[9] That assumption is consistent with the evidence reported in Fuhrer and Tootell (2008), based on the estimates of empirical Taylor rules augmented with stock price changes. Below we examine the robustness of our findings to allowing for a contemporaneous policy response to stock prices.

shock ε_t^m hitting the economy at time t are given by

$$\begin{aligned}\frac{\partial x_{t+k}}{\partial \varepsilon_t^m} &= \frac{\partial x_{t+k}}{\partial u_t'}\frac{\partial u_t}{\partial \varepsilon_t^m} \\ &= B_{t,k}S_t^{(4)} \equiv C_{t,k}\end{aligned}$$

for $k = 0, 1, 2, ...$and where $S_t^{(4)}$ denotes the fourth column of S_t. In the case of the constant coefficients model the response is just given by $\partial x_{t+k}/\partial \varepsilon_t^m = B_k S^{(4)} \equiv C_k$, where $B_k \equiv [\mathbf{A}^k]_{5,5}$.

We use Bayesian methods in order to estimate the model with time-varying coefficients. The goal of our estimation is to characterize the joint posterior distribution of the parameters of the model. To do that we use a Gibbs sampling procedure which works as follows. The parameters are divided in subsets. Parameters in each subset are drawn conditional on a particular value of the remaining parameters. The new draw is then used to draw the remaining subsets of parameters conditional on this. The procedure is repeated many times. After a burn-in period (of 50000 draws in our case) these conditional draws converge to a draw from the joint posterior.

2.1 Relation with the Existing Literature

We are not the first to analyze empirically the impact of monetary policy changes on stock prices.

Patelis (1997) analyzes the role played by monetary and financial variables in predicting stock returns. He finds that increases in the federal funds rate have a significant negative impact on predicted stock returns in the short run, but a positive one at longer horizons. That predictability works largely through the effect of federal funds rate changes on anticipated excess returns down the road, rather than dividends or expected returns.

Bernanke and Kuttner (2005) use an event-study approach, based on daily changes observed on monetary policy decision dates, to uncover the effects on stock prices of unanticipated changes in the federal funds rate. They find that a surprise 25-basis-point cut in the Federal funds rate is associated with about a 1 percent increase in stock prices. Their analysis largely attributes that response to a persistent response of the equity premium, and to a lesser extent of the relevant cash flows. They do not report, however, the dynamic response of stock prices to the monetary policy surprise. Rigobon and Sack (2004) obtain similar (but slightly larger) estimates of the response of stock prices to changes in interest rates using a heteroskedasticity-based estimator that exploits the increase in the volatility of interest rates on FOMC meeting and Humphrey-Hawkins testimony dates.

Gürkaynak et al (2005) use intraday data estimate the response of asset prices to two factors associated with FOMC decisions. The first factor corresponds, like in Bernanke and Kuttner (2005), to the unanticipated movements in the Federal funds rate target. The estimated effect on stock prices is very similar to that uncovered by Bernanke and Kuttner (2005).[10] The second factor is associated with revisions in expectations about future rates, given the funds rate target, and appears to be linked to the statement accompanying the FOMC decision. The impact of this second factor on stock prices is significant, but more muted than the first, possibly due to revisions in expectations on output and inflation which may partly offset the impact of anticipated changes in interest rates.

3 Evidence

In this section we report the impulse responses of a number of variables to a monetary policy shock, generated by our estimated VARs, both with constant and time-varying coefficients. We use quarterly U.S. time series for GDP and its deflator, the federal funds rate, and the S&P500 stock price index and the corresponding dividend series (both deflated by the GDP deflator). Our baseline sample period is 1960Q1-2011Q4. Due to the impact of the zero lower bound on the behavior of the federal funds rate since 2008 and its likely influence on our estimates we have also estimated the model ending the sample in 2007Q4 as a robustness check.

Figure 2.a displays the estimated responses to a contractionary monetary policy shock of nominal and real interest rates, (log) dividends and (log) real stock prices based on the estimated VAR with constant coefficients. The tightening of monetary policy leads to a persistent increase in both nominal and real rates, and an accompanying decline in dividends. The stock price index is also seen to decline in the short run, but it recovers fast subsequently and ends up in slightly positive territory (though the confidence bands are too large to reject long run non-neutrality). Figure 2.b displays the implied response of the "fundamental" component of the stock price, computed using (5), together with that of the observed price shown earlier. Not surprisingly, given the response of the real rate and dividends), the fundamental stock price is shown to decline sharply on impact, and to return only gradually to its initial value.

Note that (4) implies

$$\frac{\partial(q_{t+k}-q_{t+k}^F)}{\partial\varepsilon_t^m}=\gamma_{t-1}\left(\frac{\partial q_{t+k}^B}{\partial\varepsilon_t^m}-\frac{\partial q_{t+k}^F}{\partial\varepsilon_t^m}\right)$$

In the simple example of a rational bubble considered above (with exogenous dividends

[10] Similar results are obtained by D'Amico and Farka (2011) in their first-step, which involves the same intraday-data strategy as Gürkaynak et al. (2005).

and a geometric response of the real rate) we have:

$$\frac{\partial(q_{t+k}-q^F_{t+k})}{\partial\varepsilon^m_t}=\gamma_{t-1}\left(\frac{\rho^k_r}{1-\Lambda\rho_r}+\psi_t+\frac{1-\rho^k_r}{1-\rho_r}\right)$$

Thus, to the extent that a bubble is present to begin with ($\gamma_{t-1}>0$) and its contemporaneous response to the interest rate innovation is not too negative ($\psi_t \gtrsim 0$), the gap between the response of the asset price and its fundamental should be positive and increasing over time in response to a tightening of monetary policy.

As Figure 2.b makes clear the response of the gap $q_{t+k}-q^F_{t+k}$ is positive and, after one period, increasing, which points to (i) the existence of a non-negligible bubble component, and (ii) a substantial difference between the responses of the bubble and fundamental components of stock prices to a monetary policy shock.

Figures 3a-f show the evolution over time of the impulse responses to a monetary policy shock, based on our estimated VAR with time-varying coefficients. The estimated dynamic responses of nominal and real rates appear to be relatively unchanged over time, though the former shows substantially greater persistence over the last few years of the sample period (possibly due to the "distortion" created by the zero lower bound). Broadly speaking, the same holds true for the response of dividends, with the exception of a brief period in the early 1980s, when the tightening of policy appears to have a positive impact on dividends after about three years.

Our focus is, however, on the changing response of stock prices. Note that the S&P500 generally declines on impact, often substantially, in response to an exogenous monetary policy tightening. Until the late 1970s that decline is persistent, in a way consistent with the response of stock prices in the absence of a bubble. By contrast, starting in the early 1980s, the initial decline is rapidly reversed with stock prices rising quickly (and seemingly permanently) above their initial value. That phenomenon is particularly acute in the 1980s and 1990s. The previous estimated response stands in contrast with that of fundamental component, as implied by the impulse responses of the real rate and dividends, and shown in Figure 3.e Note that the pattern of the response of fundamental stock prices to a tightening of monetary policy has changed little over time, (roughly) corresponding to that obtained with the constant coefficient VAR. Figure 3.f displays the response of the gap between observed and fundamental stock prices. Note that with the exception of the early part of the sample that gap appears to be positive and growing, in a way consistent with the theory of bubbles, and in contrast with the "conventional" view. Figure 4.a provides an alternative perspective to the same evidence, by displaying the evolution over time of the impact of the monetary policy shock on the log deviations between observed and fundamental stock prices. Figure 4.b shows the estimated (bootstrap-based) probability that the same gap is positive. Note that the probability is well above 50 percent (and often much closer to unity) since the mid-80s.

Figures 5.a-5.e illustrate the changing patterns of stock price responses by showing the *average* impulse responses of both observed and fundamental prices over four alternative three-year periods: 1965Q1-1967Q4, 1976Q1-1978Q4, 1984Q4-1987Q3, and 1997Q1-1999Q4. The changing pattern of the gap between the two variables emerges clearly. The response during the first episode points to a larger drop of the observed price than in the fundamental. The evidence from the 1970s suggests a relatively similar pattern, though the observed price displays some overshooting relative to the fundamental. The estimated responses for the three-year periods before the crash of October 1987 and the burst of the dotcom bubble involve a very different pattern, with the observed price declining less than the fundamental to begin with, and then recovering faster to end up in strongly positive territory, as the theory of rational asset price bubbles would predict when a large bubble is present.

We have examined the robustness of our results to the use of earnings instead of dividends. Even though the latter is, in principle, the appropriate variable, earnings are often used in applications due to their less erratic seasonal patterns. Our findings are largely unchanged when this alternative measure of stock payoffs is used. Figure 6 illustrates that robustness by showing the dynamic response of the gap between the stock price and the fundamental to an exogenous tightening of monetary policy, using the earnings-based VAR. Note that the observed pattern of responses is very similar to that found in Figure 3.f, at least qualitatively.

4 Alternative interpretations

4.1 Time-varying equity premium

The theoretical analysis of section 1 has been conducted under the maintained assumption of risk neutrality or –equivalently, for our purposes– of a constant expected excess return (or equity premium). That assumption also underlies our definition of the fundamental component of stock prices and of the estimates of the latter's dynamic reponse to monetary policy shocks shown in the previous section. There is plenty of evidence in the literature, however, of time-varying expected excess return in stock prices, partly linked to monetary policy shocks.[11] Next we examine whether our estimated deviation between observed stock prices and the "measured" fundamental component can be plausibly interpreted as resulting from a time-varying equity premium, as an alternative to the bubble-based interpretation.

Let z_{t+1} denote the (log-linearized) excess return on stocks held between t and $t+1$,

[11] See, e.g. Thorbecke (1997), Patelis (1997).and Bekaert, Hoerova and Lo Duca (2013).

given by

$$z_{t+1} = \Lambda q_{t+1} + (1-\Lambda)d_{t+1} - q_t - r_t$$

In the absence of a bubble, we can write the equilibrium stock price

$$q_t = const + \sum_{k=0}^{\infty} \Lambda^k \left[(1-\Lambda)E_t\{d_{t+k+1}\} - E_t\{r_{t+k}\} - E_t\{z_{t+k+1}\}\right]$$

Thus, the dynamic response of the stock price to an exogenous monetary policy shock is given by

$$\frac{\partial q_{t+k}}{\partial \varepsilon_t^m} = \sum_{j=0}^{\infty} \Lambda^j \left((1-\Lambda)\frac{\partial d_{t+k+j+1}}{\partial \varepsilon_t^m} - \frac{\partial r_{t+k+j}}{\partial \varepsilon_t^m} - \frac{\partial z_{t+k+j+1}}{\partial \varepsilon_t^m} \right)$$

Then it follows that the gap between the response of the observed price and the response of the fundamental component computed under the assumption of risk neutrality are related to the equity premium response according to the equation:

$$\frac{\partial q_{t+k}}{\partial \varepsilon_t^m} - \frac{\partial q_{t+k}^F}{\partial \varepsilon_t^m} = -\sum_{j=0}^{\infty} \Lambda^j \frac{\partial z_{t+k+j+1}}{\partial \varepsilon_t^m}$$

for $k = 0, 1, 2, ...$and where as above $\frac{\partial q_{t+k}^F}{\partial \varepsilon_t^m} \equiv \sum_{j=0}^{\infty} \Lambda^j \left((1-\Lambda)\frac{\partial d_{t+k+j+1}}{\partial \varepsilon_t^m} - \frac{\partial r_{t+k+j}}{\partial \varepsilon_t^m} \right)$ is the fundamental stock price under risk neutrality.

Thus, an interpretation of the evidence above that abstracts from the possibility of bubbles and relies instead on a time-varying equity premium requires that the latter declines substantially and persistently in response to a tightening of monetary conditions. That implication is at odds with the existing evidence on the response of excess stock returns (e.g. Patelis (1997), Bernanke and Kuttner (2005)) or variables that should be closely related to it, like the VIX (Bekaert et al. (2013)).

4.2 Simultaneity

The estimates reported above were obtained under the identifying assumption that the Federal Reserve did not respond contemporaneously (i.e. within the quarter) to stock price innovations. That assumption is consistent with the "pre-crisis consensus" according to which central banks should focus exclusively on stabilizing inflation and the output gap. It is also consistent with formal evidence in Fuhrer and Tootell (2008) based on estimated interest rate rules using real time Greenbook forecasts, though that evidence does not rule out the possibility of an indirect response to stock prices, based on their potential ability to predict output or inflation developments.

Here we examine the robustness of our findings to relaxing that constraint, by allowing for some (contemporaneous) simultaneity in the determination of interest rates and stock prices. More specifically, we re-estimate our empirical model under the assumption that current log change in stock prices enter the interest rate rule with a coefficient 0.02. This implies that, ceteris paribus, a ten percentage point increase in stock prices within a quarter triggers a 20 basis points rise in the federal funds rate. The previous assumption is consistent with the estimated reaction of monetary policy to the stock market changes obtained by Rigobon and Sack (2003) using an approach that exploits heteroskedasticity in stock price shocks to identify the coefficient measuring that reaction.[12]

The estimated responses of interest rates and dividends to a monetary policy shock (not shown) are hardly affected by the use of this alternative identification scheme. But the same cannot be said for stock prices: the latter now decline persistently throughout the sample period in response to a tightening of monetary policy, as shown in Figure 7. Furthermore, and most importantly for our purposes, the gap between the observed price and the estimated fundamental price also declines strongly in response to the same shock, as shown in Figure 8. The latter response is consistent, at least in a qualitative sense, with the conventional wisdom regarding the impact of monetary policy on stock price bubbles, and contrast starkly with the evidence based on our baseline specification.

If one accepts this alternative identifying assumption as correct, the findings obtained in the previous section should be interpreted as spurious, and driven by biased estimates of matrices $\{S_t\}$ resulting from the imposition of an incorrect identifying assumption. Figure 9 displays the evolution over time of the stock price response to the tightening of monetary policy, both on impact and after four quarters, for four alternave calibrations of the contemporaneous stock price coefficient in the interest rate rule: 0.0, 0.01, 0.02, and 0.03. We see that estimates of the effects of monetary policy on stock prices are rather sensitive to the calibration of that parameter. Those differences emerge at all horizons, even on impact. In a nutshell, the larger is the calibrated stock price coefficient in the interest rate rule, the smaller (i.e. more negative) is the estimated effect of an interest rate shock on stock prices. That negative conditional comovement is required in order to compensate for the strong positive comovement that arises as a result of non-monetary policy shocks, due to the endogenous policy response to stock price movements embedded in the rule.

The previous interpretation, however, is subject to an important caveat. In a recent paper, Furlanetto (2011) has revisited the evidence of Rigobon and Sack (2003) using data that extends over a longer sample period (1988-2007) and focusing on the stability over time in the estimates of the monetary policy response to stock prices.[13] He shows that the

[12] D'Amico and Farka (2011) use an alternative two-step procedure to identify the policy response to stock prices, obtaining a similar estimate of the response coefficient (about 0.02). Furlanetto (2011) revisits de Rigobon-Sack evidence and concludes that the positive estimated reaction is largely driven by the Fed response to the stock market crash of 1987.

[13] In addition, he also examines the evidence for six other economies (Australia, Canada, New Zealand,

main finding in Rigobon and Sack (2003) is largely driven by a single episode: the Fed's interest rate cuts in response to the stock market crash in 1987. When the same empirical model is re-estimated using post-1988 data, the estimated policy response is much smaller or insignificant. The Furlanetto evidence has an important implication for the present paper, for it it suggests that our baseline specification may have been a good approximation, possibly with the exception of the period around 1987. Given that our empirical framework allows the model's coefficients to vary over time, that "transitory" misspecification should not distort the estimated responses for other "segments" of the sample. Thus, and conditional on Furlanetto's findings, our evidence pointing to an eventual positive (and growing) response of stock prices (in both levels and deviations from fundamentals) to a tightening of monetary policy should be viewed as valid., while the estimates using the alternative specification would likely be distorted by the imposition of an identifying assumption that is invalid for much of the sample.

5 Concluding Remarks

Proposals for a "leaning against the wind" monetary policy in response to perceived deviations of asset prices from fundamentals rely on the assumption that increases in interest rates will succeed in shrinking the size of an emerging asset price bubble. Yet, and despite the growing popularity of such proposals, no evidence seems to be available providing support for that link.

In the present paper we have provided evidence on the response of stock prices to monetary policy shocks, and tried to use that evidence to evaluate the empirical merits of the "conventional" view according to which the size of the bubble component of stock prices should decline in response to an exogenous increase in interest rates.

Our evidence is based on an estimated vector-autoregression with time-varying coefficients, applied to quarterly US data for GDP, inflation, dividends, the federal funds rate, and a stock price index (S&P500). Under our baseline specification, which assumes no contemporaneous response of monetary policy to asset prices, the evidence points to protracted episodes in which stock prices increase persistently in response to an exogenous tightening of monetary policy. That response is clearly at odds with the "conventional" view on the effects of monetary policy on bubbles, as well as with the predictions of bubbleless models. We also argue that it is unlikely that such evidence be accounted for by an endogenous response of the equity premium to the monetary policy shocks.

The previous findings are overturned when we impose a contemporaneous interest rate response to stock prices consistent with the evidence in Rigobon and Sack (2003): under

Norway, Sweden and the United Kingdom). He finds evidence of a significant endogenous response to stock prices only in Australia.

this alternative specification our evidence points to a decline in stock prices in response to a tightening of monetary policy, beyond that warranted by the estimated reponse of the fundamental price. Recent independent evidence by Furlanetto (2011), however, calls into question the relevance of this alternative specification.

Further research seems to be needed to improve our understanding of the effect of interest rate changes on asset price bubbles. That understanding is a necessary condition before one starts thinking about how monetary policy *should* respond to asset prices. We hope to have contributed to that task by providing some evidence that calls into question the prevailing dogma among advocates of "leaning against the wind" policies, namely, that a rise in interest rates will help disinflate an emerging bubble.

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Appendix

We use Bayesian methods to estimate our empirical model with time-varying coefficients. To draw from the joint posterior distribution of model parameters we use the Gibbs sampling algorithm along the lines described in Primiceri (2005). The algorithm draws sets of coefficients from known conditional posterior distributions. Under some regularity conditions, the draws converge to a draw from the joint posterior after a burn-in period. Let T be the length of the sample and let z be a generic $(q \times 1)$ vector. We denote z^T the sequence $[z_1', ..., z_T']'$. Each repetition of the algorithm is composed of seven steps where a draw for a set of parameters is made conditional on the value of the remaining parameters. Below we report the conditional distributions used in each of the steps:[14]

1. $p(\sigma^T | x^T, \theta^T, \phi^T, \Omega, \Xi, \Psi, s^T)$
2. $p(\phi^T | x^T, \theta^T, \sigma^T, \Omega, \Xi, \Psi, s^T)$
3. $p(\theta^T | x^T, \sigma^T, \phi^T, \Omega, \Xi, \Psi, s^T)$
4. $p(s^T | x^T, \theta^T, \sigma^T, \phi^T, \Omega, \Xi, \Psi)$[15]
5. $p(\Omega | x^T, \theta^T, \sigma^T, \phi^T, \Xi, \Psi, s^T)$
6. $p(\Xi | x^T, \theta^T, \sigma^T, \phi^T, \Omega, \Psi, s^T)$
7. $p(\Psi_i | x^T, \theta^T, \sigma^T, \phi^T, \Omega, \Xi, s^T)$, $i = 1, 2, 3, 4$

Priors Specification

We make the following assumptions about the priors densities. First, the hyperparameters Ω, Ξ and Ψ and the initial states are independent. Second, the priors for the initial states θ_0, ϕ_0 and $\log \sigma_0$ are assumed to be normally distributed. Third, the priors for the hyperparameters, Ω, Ξ and Ψ are assumed to be distributed as independent inverse-Wishart. More precisely

$$\begin{aligned} P(\theta_0) &= N(\hat{\theta}, \hat{V}_\theta) \\ P(\log \sigma_0) &= N(\log \hat{\sigma}_0, I_n) \\ P(\phi_{i0}) &= N(\hat{\phi}_i, \hat{V}_{\phi_i}) \\ P(\Omega) &= IW(\Omega_0^{-1}, \rho_1^0) \\ P(\Xi) &= IW(\Xi_0^{-1}, \rho_2^0) \\ P(\Psi_i) &= IW(\Psi_{0i}^{-1}, \rho_{3i}^0) \end{aligned}$$

where ρ_i^0 $(i = 1, 2, 3)$ refers to the prior degrees of freedom and the scale matrices are parameterized as follows $\Omega_0 = \lambda_1 \rho_1 \hat{V}_\theta$, $\Xi_0 = \lambda_2 \rho_2 I_n$ and $\Psi_{0i} = \lambda_{3i} \rho_{3i} \hat{V}_{\phi_i}$.

[14] Notice that the following ordering of the seven steps is not subject to the problem discussed in Del Negro and Primiceri (2012) since the draw of s^T is made after the draw id σ^T and θ^T

[15] See below the definition of s^T.

Prior means and variances are calibrated using a time invariant VAR for x_t estimated using the first $\tau = 48$ observations. $\hat{\theta}$ and $\hat{V}_\theta$ are set equal to the OLS estimates. Let $\hat{u}_{i+1,t}$ be the residual of the $i+1$-th equation of the initial time-invariant VAR. We apply the same decomposition discussed in the text to the VAR residuals covariance matrix, $\hat{\Sigma} = \hat{F}\hat{D}\hat{F}'$. $\log \hat{\sigma}_0$ is set equal to the log of the diagonal elements of $\hat{D}^{1/2}$. $\hat{\phi}_i$ is set equal to the OLS estimates of the coefficients of the regression of $\hat{u}_{i+1,t}$ on $-\hat{u}_{1,t}, ..., -\hat{u}_{i,t}$ and $\hat{V}_{\phi_i}$ equal to estimated variances.

The degrees of freedom for the priors on the covariance matrices ρ_1 and ρ_2 are set equal to the number of rows Ω_0^{-1} and I_n plus one respectively while ρ_{3i} is $i+1$ for $i = 1, ..., n-1$. Finally we assume $\lambda_1 = 0.001$, $\lambda_2 = 0.01$ and $\lambda_3 = 0.01$.

Gibbs sampling algorithm

To draw realizations from the posterior distribution we use T observations from $\tau/2+1$ up to $\bar{T}$.[16] The algorithm works as follows:

Step 1: sample from $p(\sigma^T | x^T, \theta^T, \phi^T, \Omega, \Xi, \Psi, s^T)$. To draw σ^T we use the algorithm of Kim, Shephard and Chibb (1998, KSC hereafter). Consider the system of equations $x_t^* \equiv F_t^{-1}(x_t - Z_t'\theta_t) = D_t^{1/2} u_t$, where $u_t \sim N(0, I)$, $Z_t = (I_n \otimes z_t)$, and $z_t = [1_n, x_{t-1} ... x_{t-p}]'$. Conditional on x^T, θ^T, and ϕ^T, x_t^* is observable. Squaring and taking logs, we obtain

$$x_t^{**} = 2r_t + \upsilon_t \tag{17}$$

$$r_t = r_{t-1} + \xi_t \tag{18}$$

where $x_{i,t}^{**} = \log({x_{i,t}^*}^2)$[17], $\upsilon_{i,t} = \log(u_{i,t}^2)$ and $r_t = \log \sigma_{i,t}$. Since, the innovation in (17) is distributed as $\log \chi^2(1)$, we follow KSC and we use a mixture of 7 normal densities with component probabilities q_j, means $m_j - 1.2704$, and variances v_j^2 (j=1,...,7) to transform the system in a Gaussian one, where $\{q_j, m_j, v_j^2\}$ are chosen to match the moments of the $\log \chi^2(1)$ distribution (see Table A1 for the values used).

j	q_j	m_j	v_j^2
1.0000	0.0073	-10.1300	5.7960
2.0000	0.1056	-3.9728	2.6137
3.0000	0.0000	-8.5669	5.1795
4.0000	0.0440	2.7779	0.1674
5.0000	0.3400	0.6194	0.6401
6.0000	0.2457	1.7952	0.3402
7.0000	0.2575	-1.0882	1.2626

[16] We start the sample from $\tau/2+1$ instead of τ in order to not to lose too many data points.

[17] We do not use any offsetting constant sine given that the variables are in logs times 100 we do not have numerical problems.

Let s_t be a vector whose elements indicate the element of the mixture to be used for each element of υ_t at time t. Conditional on s^T, $(\upsilon_{i,t}|s_{i,t} = j) \sim N(m_j - 1.2704, v_j^2)$. Therefore we can use the algorithm of Carter and Khon (1994, CK henceforth) to draw r_t from $N(r_{t|t+1}, R_{t|t+1})$, where $r_{t|t+1} = E(r_t|r_{t+1}, x^t, \theta^T, \phi^T, \Omega, \Xi, \Psi, s^T,)$ and $R_{t|t+1} = Var(r_t|r_{t+1}, x^t, \theta^T, \phi^T, \Omega, \Xi, \Psi, s^T)$ are the conditional mean and variance obtained from the backward recursion equations.

Step 2: sample from $p(\phi^T|x^T, \theta^T, \sigma^T, \Omega, \Xi, \Psi, s^T)$. Consider again $F_t^{-1}(x_t - X_t'\theta_t) = F_t^{-1}\hat{x}_t = D_t^{1/2}u_t$. Conditional on θ^T, $\hat{x}_t$ is observable. Each equation of the above system can be written as

$$\hat{x}_{1,t} = \sigma_{1,t}u_{1,t} \tag{19}$$

$$\hat{x}_{i,t} = -\hat{x}_{[1,i-1],t}\phi_{i,t} + \sigma_{i,t}u_{i,t} \quad i = 2, ..., n \tag{20}$$

where $\sigma_{i,t}$ and $u_{i,t}$ are the ith elements of σ_t and u_t respectively and $\hat{x}_{[1,i-1],t} = [\hat{x}_{1,t}, ..., \hat{x}_{i-1,t}]$. Under the block diagonality of Ψ, the algorithm of CK can be applied equation by equation to draw $\phi_{i,t}$ from a $N(\phi_{i,t|t+1}, \Phi_{i,t|t+1})$, where $\phi_{i,t|t+1} = E(\phi_{i,t}|\phi_{i,t+1}, x^t, \theta^T, \sigma^T, \Omega, \Xi, \Psi)$ and $\Phi_{i,t|t+1} = Var(\phi_{i,t}|\phi_{i,t+1}, x^t, \theta^T, \sigma^T, \Omega, \Xi, \Psi)$.

Step 3: sample from $p(\theta^T|x^T, \sigma^T, \phi^T, \Omega, \Xi, \Psi, s^T)$.[18] Conditional on all other parameters and the observables we have

$$x_t = Z_t'\theta_t + \varepsilon_t \tag{21}$$

$$\theta_t = \theta_{t-1} + \omega_t \tag{22}$$

θ_t is drawn from a $N(\theta_{t|t+1}, P_{t|t+1})$, where $\theta_{t|t+1} = E(\theta_t|\theta_{t+1}, x^t, \sigma^T, \phi^T, \Omega, \Xi, \Psi)$ and $P_{t|t+1} = Var(\theta_t|\theta_{t+1}, x^t, \sigma^T, \phi^T, \Omega, \Xi, \Psi)$ are obtained using CK.

Step 4: sample from $p(s^T|x^T, \theta^T, \sigma^T, \phi^T, \Omega, \Xi, \Psi)$. Conditional on $y_{i,t}^{**}$ and r^T, we independently sample each $s_{i,t}$ from the discrete density defined by $Pr(s_{i,t} = j|x_{i,t}^{**}, r_{i,t}) \propto f_N(y_{i,t}^{**}|2r_{i,t} + m_j - 1.2704, v_j^2)$, where $f_N(y|\mu, \sigma^2)$ denotes a normal density with mean μ and variance σ^2.

Step 5: sample from $p(\Omega|x^T, \theta^T, \sigma^T, \phi^T, \Xi, \Psi, s^T)$. Conditional on the other coefficients and the data, Ω has an Inverse-Wishart posterior density with scale matrix $\Omega_1^{-1} = (\Omega_0 + \sum_{t=1}^T \Delta\theta_t(\Delta\theta_t)')^{-1}$ and degrees of freedom $\rho_1^1 = \rho_1^0 + T$, where Ω_0^{-1} where T is the number of observation used for estimation. To draw a realization for Ω make ρ_1^1 independent draws z_i (i=1,...,ρ_1^1) from $N(0, \Omega_1^{-1})$ and compute $\Omega = (\sum_{i=1}^{\rho_1^1} z_i z_i')^{-1}$ (see Gelman et. al., 1995).

[18] Sampling from $p(\theta^T|x^T, \sigma^T, \phi^T, \Omega, \Xi, \Psi, s^T)$ is equivalent to sampling from $p(\theta^T|x^T, \sigma^T, \phi^T, \Omega, \Xi, \Psi)$ since σ^T has been drawn conditioning on the new value of s^T. This makes the model a linear state space and standard techniques can be applied.

Step 6: sample from $p(\Xi|x^T, \theta^T, \sigma^T, \phi^T, \Omega, \Psi, s^T)$. Conditional the other coefficients and the data, Ξ has an Inverse-Wishart posterior density with scale matrix $\Xi_1^{-1} = (\Xi_0 + \sum_{t=1}^{T} \Delta \log \sigma_t (\Delta \log \sigma_t)')^{-1}$ and degrees of freedom $\rho_2^1 = \rho_2^0 + T$ Draws are obtained as in step 5.

Step 7: sample from $p(\Psi_i|x^T, \theta^T, \sigma^T, \phi^T, \Omega, \Xi, s^T)$ for $i = 1, 2, 3, 4$. Conditional on the other coefficients and the data, Ψ_i has an Inverse-Wishart posterior density with scale matrix $\Psi_{i,1}^{-1} = (\Psi_{i,0} + \sum_{t=1}^{T} \Delta\phi_{i,t}(\Delta\phi_{i,t})')^{-1}$ and degrees of freedom $\rho_{3i}^1 = \rho_{3i}^0 + T$ Draws are obtained as in step 5 for all $i = 1, 2, 3, 4$.

We make 52000 draws discarding the first 50000 and collecting one out of two of the remaining 2000 draws. Parameters convergence is assessed using trace plots.

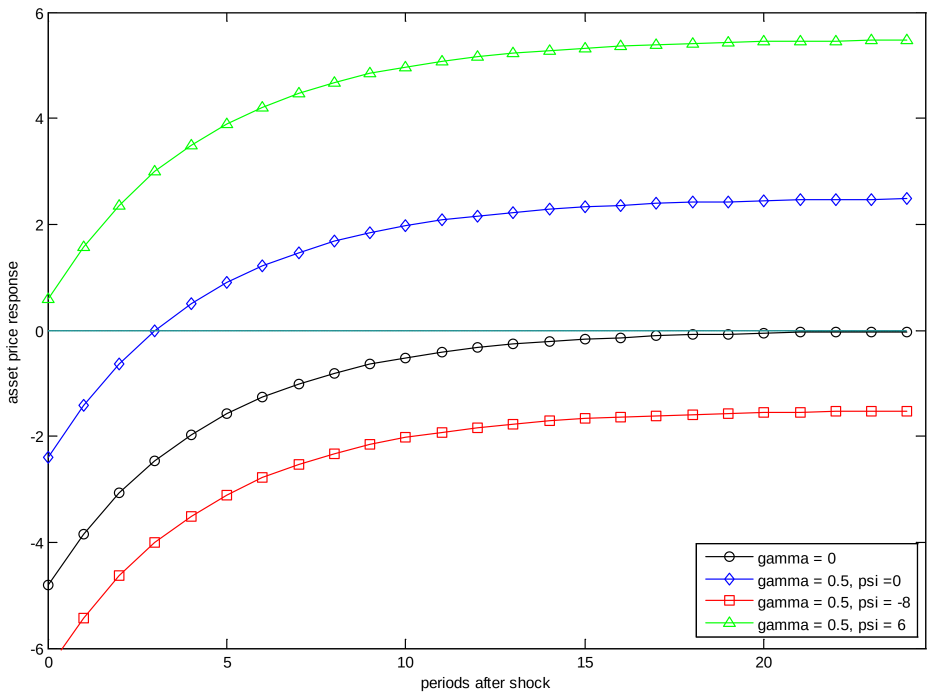
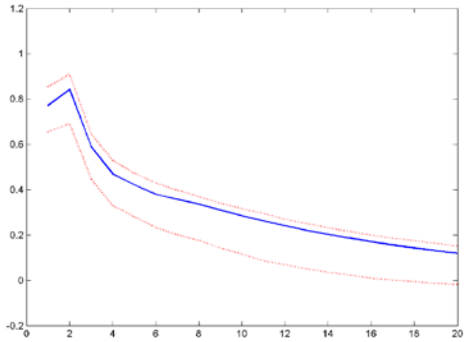
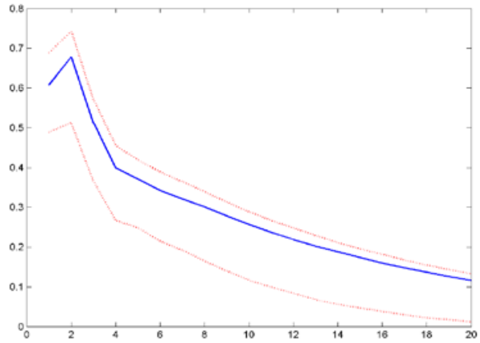


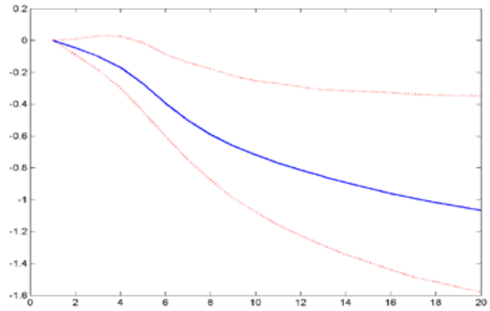
Figure 1 : Asset Price Response to an Exogenous Interest Rate Increase: Alternative Calibrations



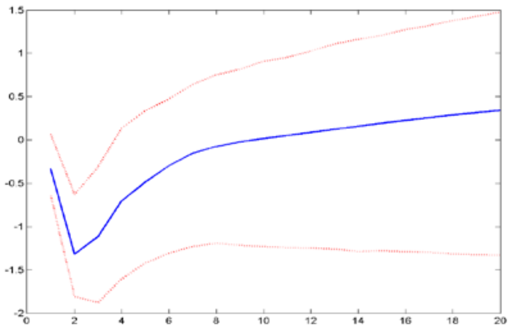
Nominal interest rate



Real interest rate



Dividends



Stock prices

Figure 2.a : Estimated Responses to Monetary Policy Shock

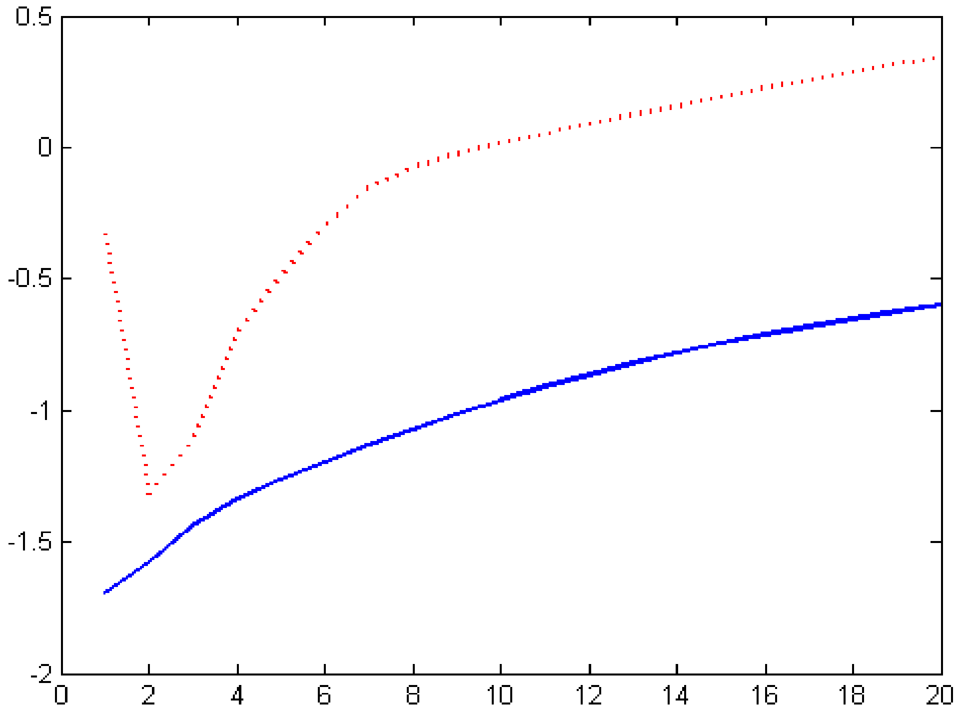


Figure 2.b : Estimated Responses to Monetary Policy Shock

Observed (red, dotted) vs. Fundamental (blue, solid) Stock Price

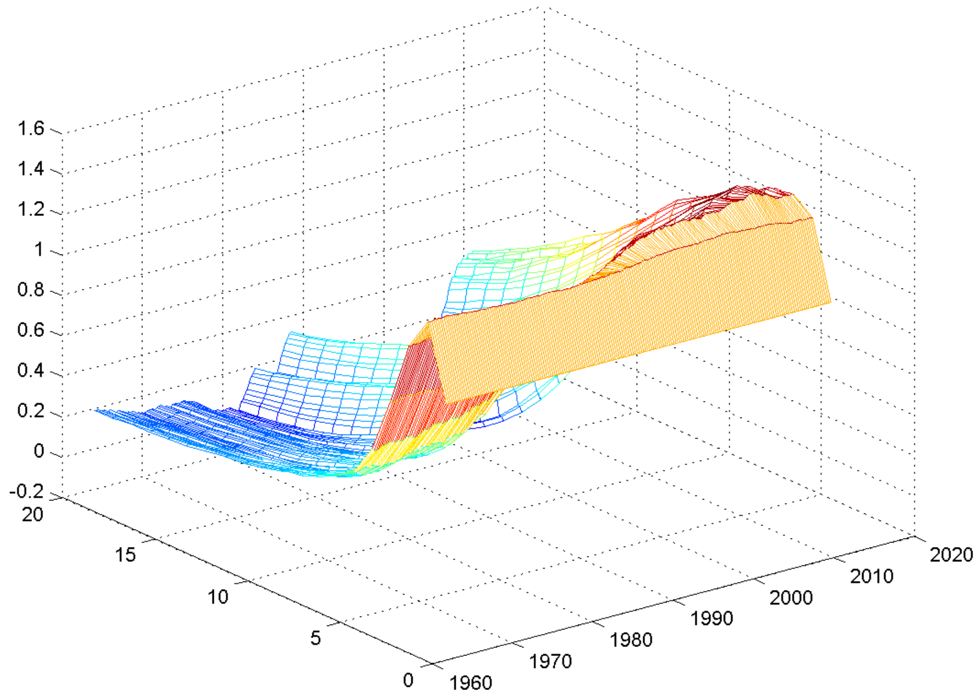


Figure 3.a : Estimated Responses to Monetary Policy Shock: TVC-VAR
Nominal Interest Rate

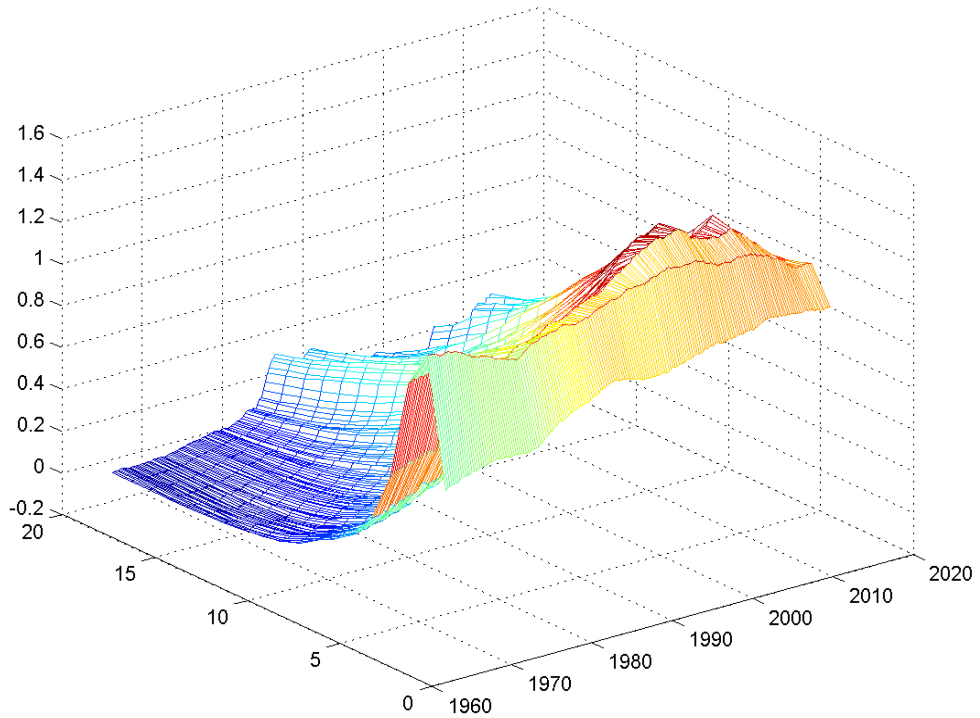


Figure 3.b : Estimated Responses to Monetary Policy Shock: TVC-VAR
Real Interest Rate

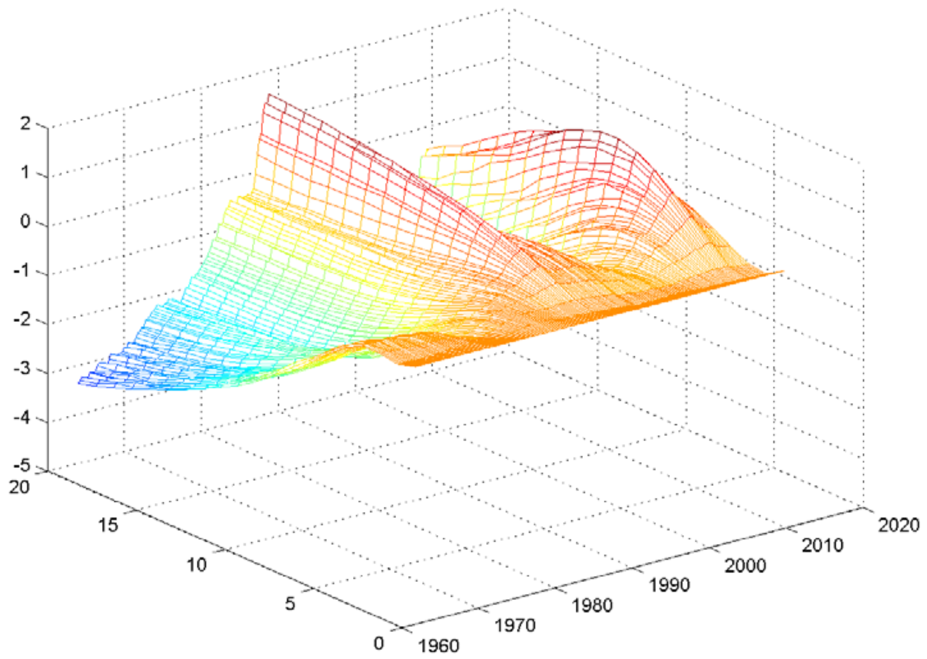


Figure 3.c : Estimated Responses to Monetary Policy Shock: TVC-VAR
Dividends

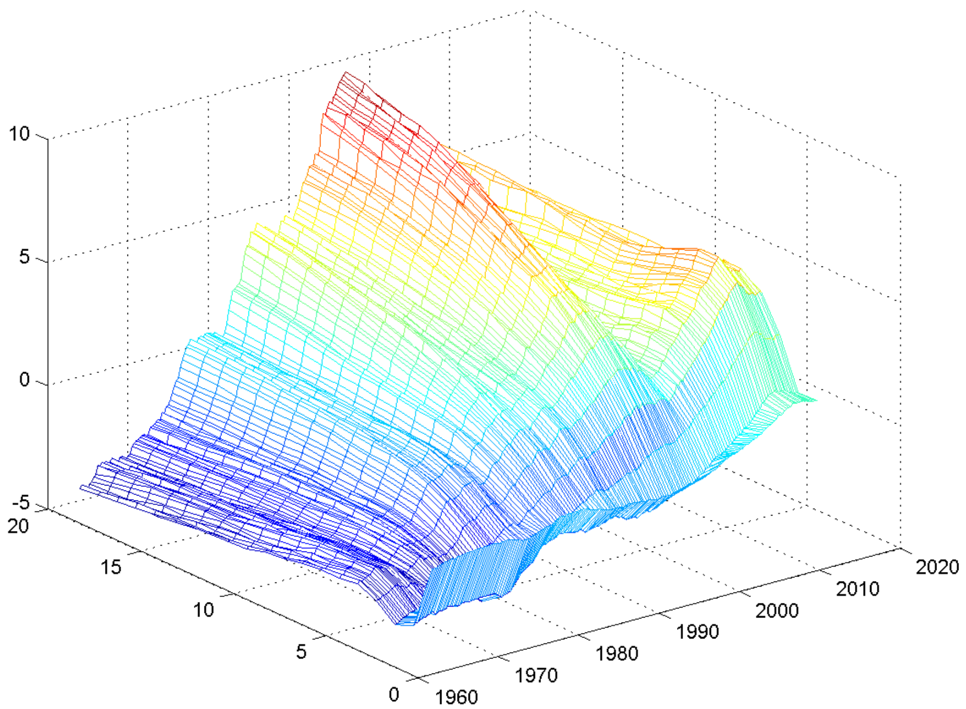


Figure 3.d : Estimated Responses to Monetary Policy Shock: TVC-VAR
Stock Prices

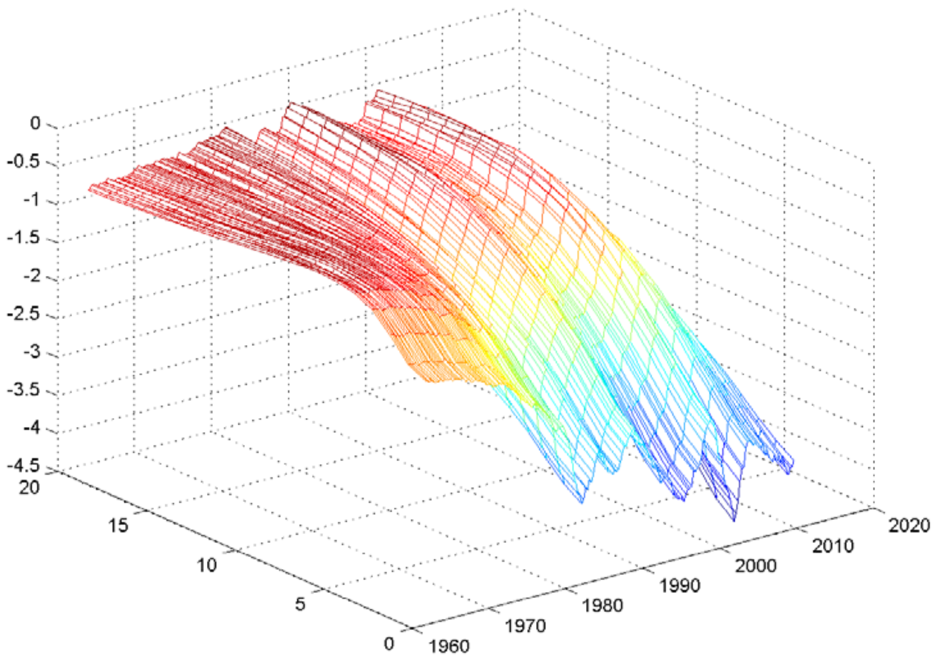


Figure 3.e : Estimated Responses to Monetary Policy Shock: TVC-VAR
Fundamental Stock Price

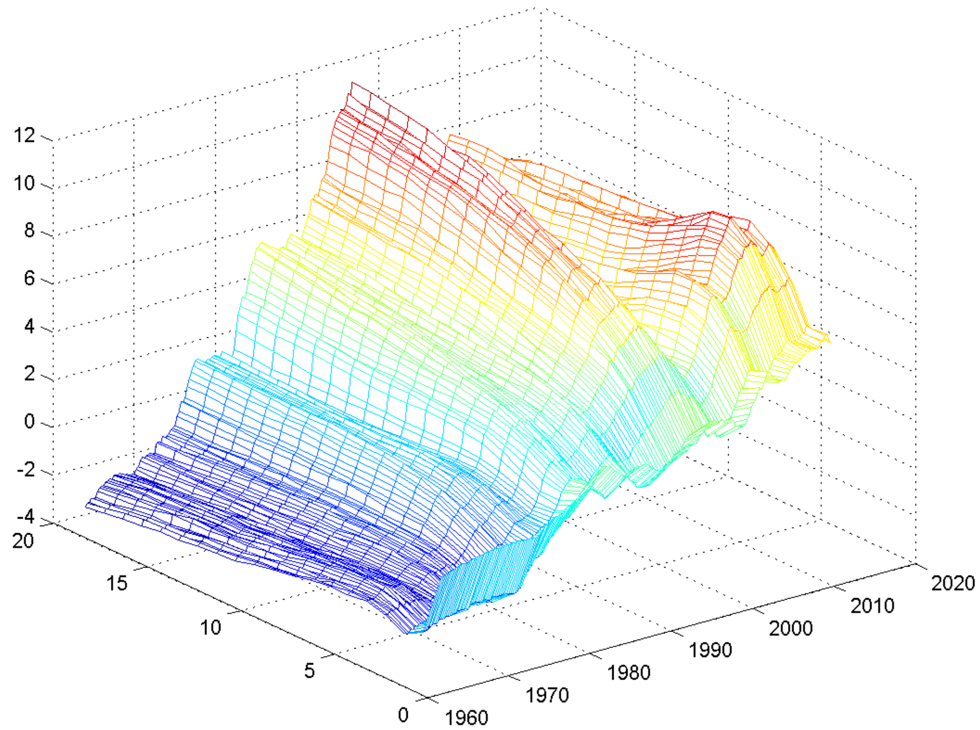


Figure 3.f : Estimated Responses to Monetary Policy Shock: TVC-VAR
Observed minus Fundamental Stock Price

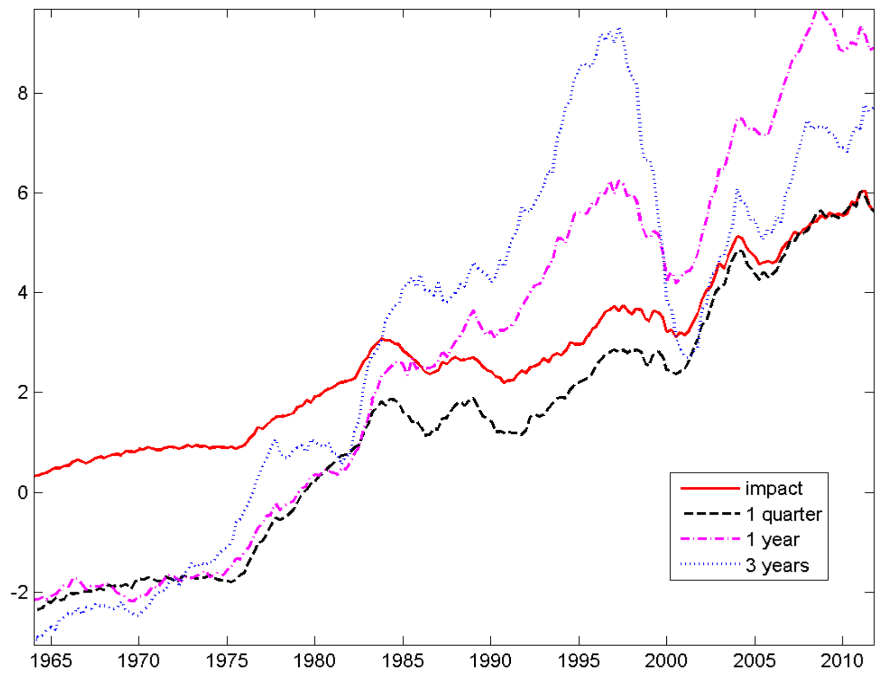


Figure 4.a : Response of $q - q^F$ at different horizons

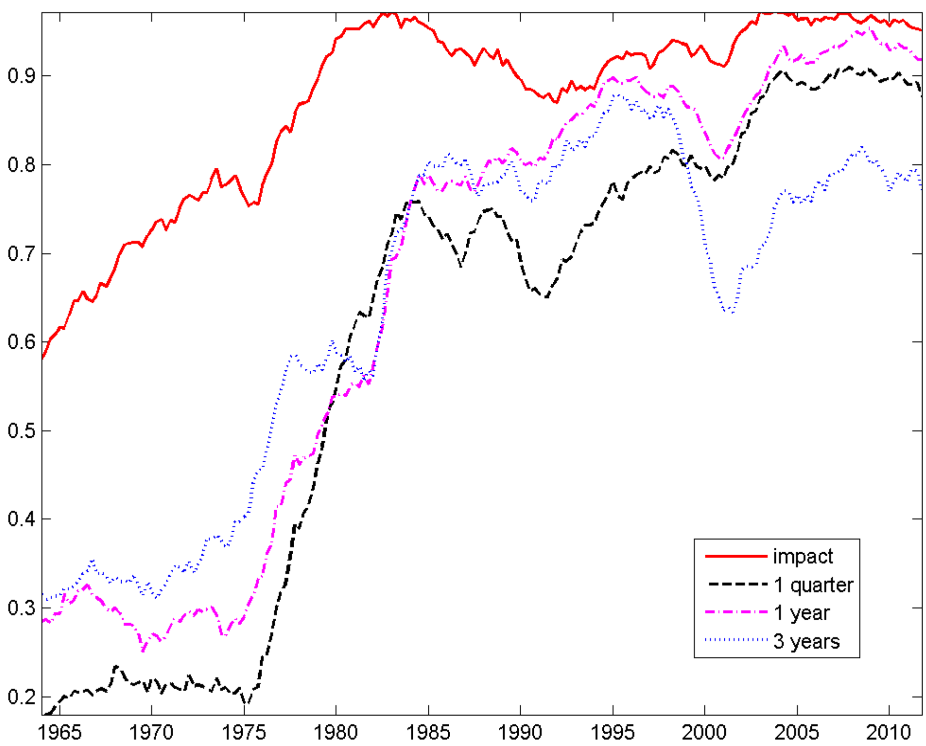
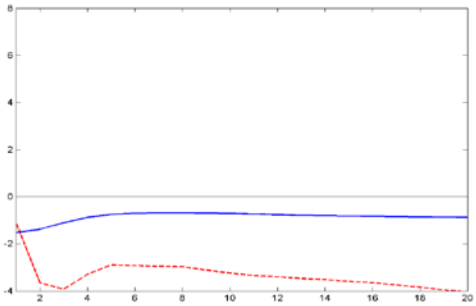
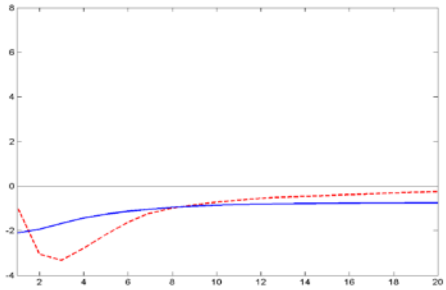


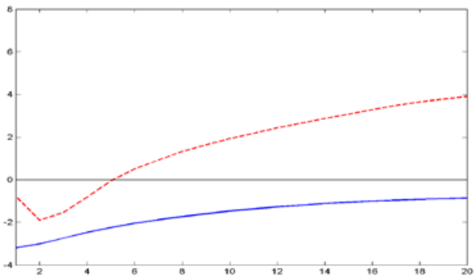
Figure 4.b : Probability of a positive response of $q - q^F$ at different horizons



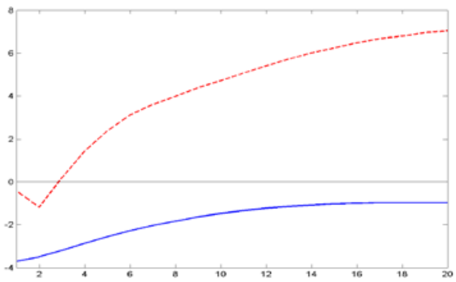
a) 1965Q1-1967Q4



b) 1976Q1-1978Q4



c) 1984Q4-1987Q3



d) 1997Q1-1999Q4

Figure 5: Estimated Responses to Monetary Policy Shock: Episode Analysis
Observed (red) vs. Fundamental (blue) Stock Prices

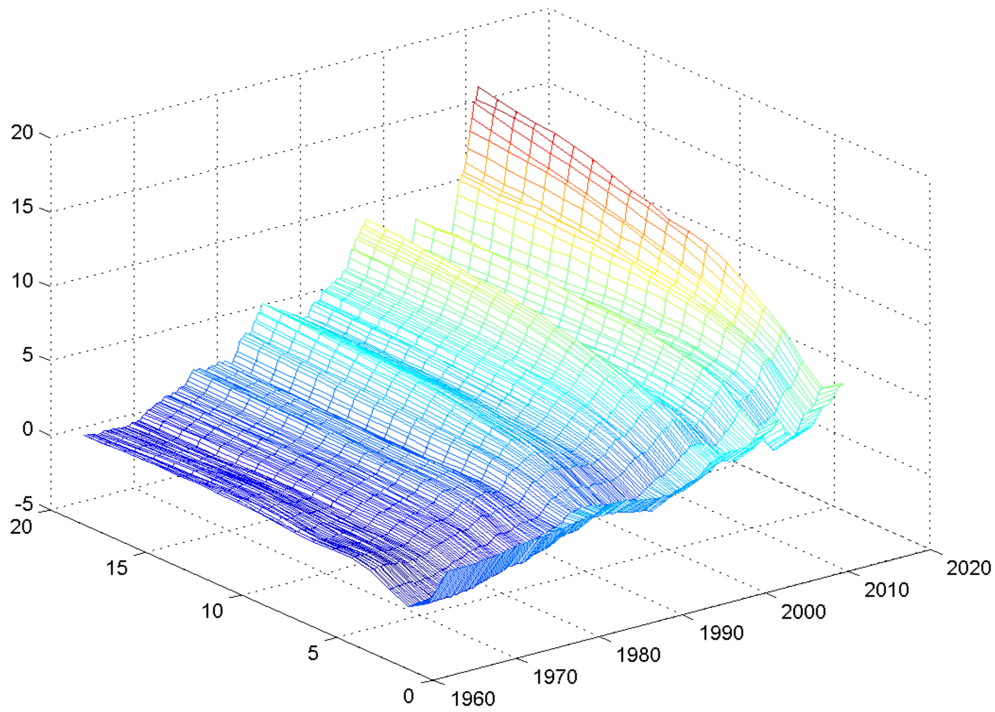


Figure 6: Estimated Responses to Monetary Policy Shock: Earnings-based
Observed minus Fundamental Stock Price

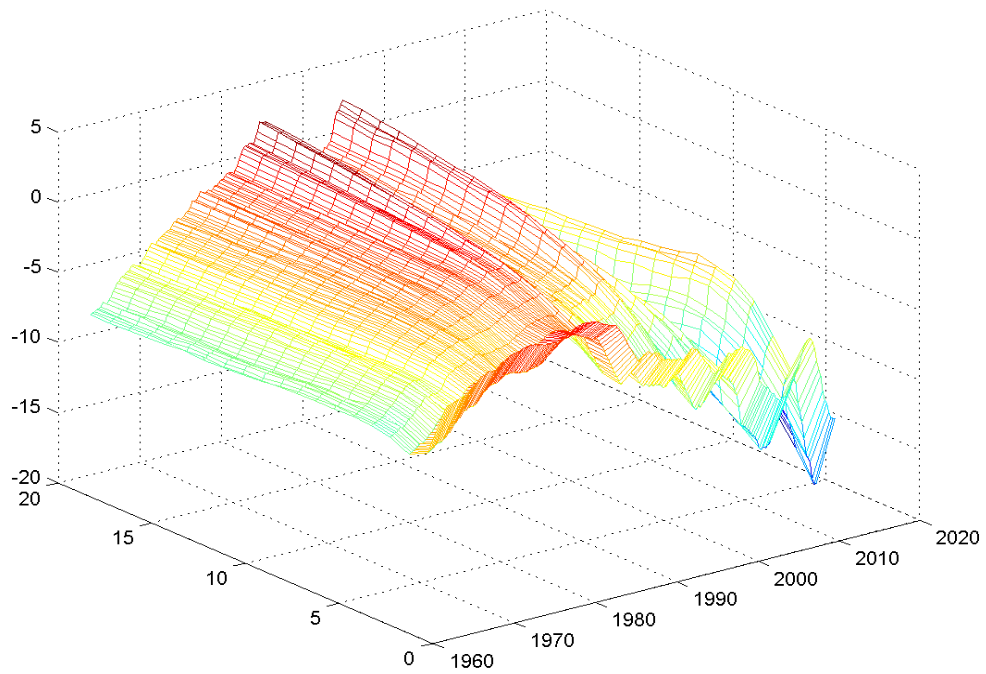


Figure 7: Estimated Responses to Monetary Policy Shock: RS calibration
Stock Prices

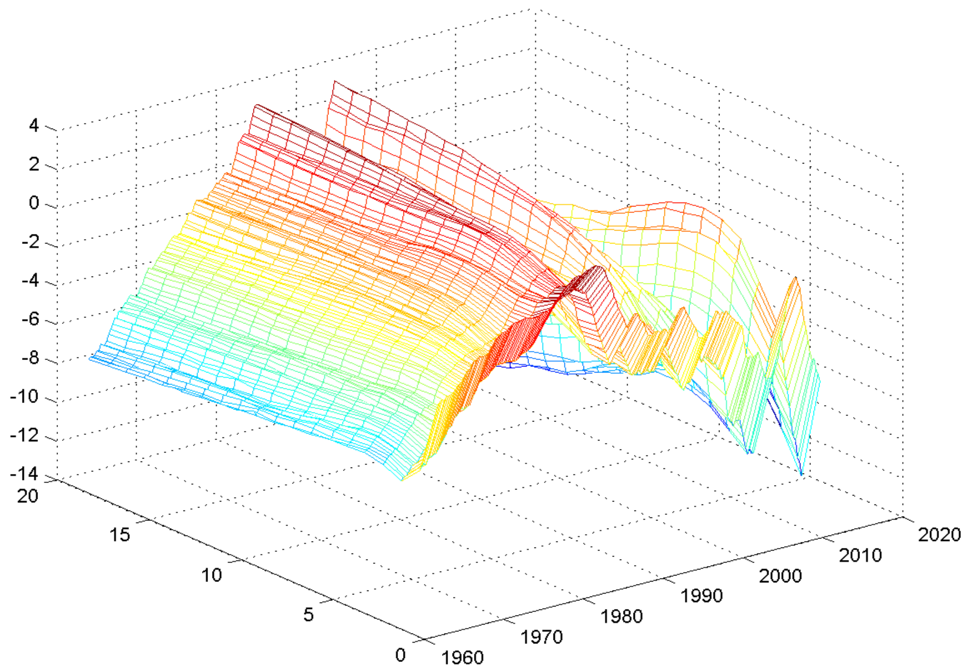
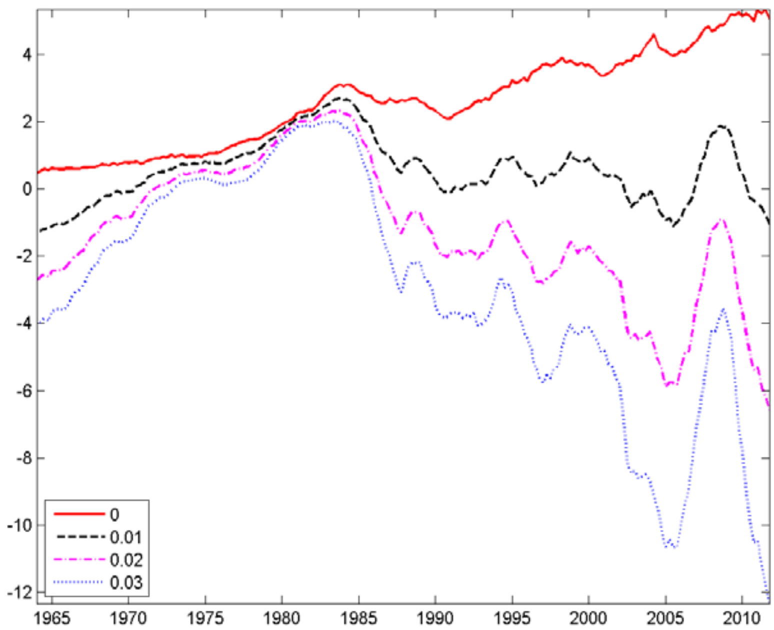
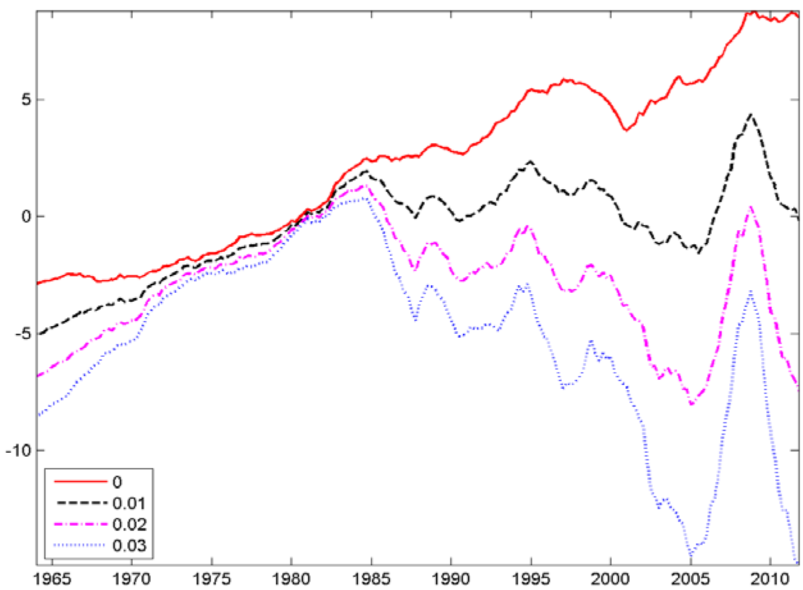


Figure 8: Estimated Responses to Monetary Policy Shock: RS calibration
Observed minus Fundamental Stock Price



Response on impact



Response after 4 quarters

Figure 9 : Estimated Responses to Monetary Policy Shock: Sensitivity Analysis
Observed minus Fundamental Stock Price