ESSAYS ON THE STATE DEPENDENT EFFECTS OF MONETARY POLICY AND FISCAL POLICY

A Dissertation

by

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Submitted to the Office of Graduate and Professional Studies of
Texas A&M University
in partial fulfillment of the requirements for the degree of

DOCTOR OF PHILOSOPHY

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August 2015

Major Subject: Economics

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ABSTRACT

This dissertation analyzes the effects of monetary policy and fiscal policy from a state-dependent perspective. The first chapter is on the dynamic effect of monetary policy on asset price. Employing a two-state threshold local projection method, we find that when the Fed increases the Federal Funds rate, the stock price decreases in normal times, but increases during bubbly episodes. We allow time-varying risk premium and show that this result is driven by both the asymmetric effects on fundamentals and the existence of bubbles. Moreover, the paper captures the effect of an exogenous tightening monetary shock on stock prices as an increasing function of the size of bubbles, using a flexible semiparametric varying-coefficient model specification. The state-dependent evidence is more informative in measuring monetary policy effects than linear or time-varying methods, and is also robust to different identification schemes and various definitions of bubbles. This paper points out two important transmission channels of monetary policy on asset price: risk premium and asset bubbles, which are often ignored in theoretical models. On the policy side, our empirical analysis suggests that central banks should be cautious about adopting "leaning against bubble" monetary policies when the bubble size is relatively large. Another contribution is that we propose a novel empirical framework to study generalized state-dependent impulse response functions, a methodology which should have many applications in macroeconomics. The second chapter uses more than one hundred years of US historical data to examine the fiscal multiplier and how it may differ during different economic conditions. Using the flexible semiparametric varying coefficient method in the framework of local projections, we directly model the fiscal multiplier as a function of various state variables. The paper shows that

the U.S. fiscal multiplier is slightly below one and approximately the same, during periods of slack as compared to normal times. Our results suggest that fiscal policy was not necessarily a more powerful tool to stimulate aggregate demand during the "Great Recession".

DEDICATION

I would like to dedicate this Doctoral dissertation to my family.

ACKNOWLEDGEMENTS

I would like to acknowledge the inspirational instruction and guidance of my advisor, Dr. Dennis W. Jansen, you have been a tremendous mentor for me. I would like to thank you for encouraging my research and for allowing me to grow as an economist. Your advice on both research as well as on my career have been priceless. I would also like to thank Dr. Sarah Zubairy, who has given me a deep appreciation and love for the beauty and detail of macroeconomics and guided me to study the state-dependent effects of macroeconomic policies. My other committee members, Dr. Yuzhe Zhang and Dr. Yong Chen, also helped me a lot on how to become a good researcher in the subject of economics and finance. I take this opportunity to express gratitute to all of the Department faculty members for their help and support.

Moreover, I thank my fellow Ph. D classmates in the Economics Department: Won Gi Kim, Hongjun Li, Wei Long, Guannan Liu, Yutang Shi, for the stimulating discussions, and for all the fun we have had in the last five years.

Finally, I also place on record, my sense of gratitude to one and all, who directly or indirectly, have lemt their hand in this venture.

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1. INTRODUCTION TO RESEARCH

This dissertation includes two empirical works on the state-dependent effects of macroeconomic policies. They are unrelated in the subject matter, but they share several similarities from the methodology perspective. The first work investigate the effects of interest rate monetary policy on asset price during normal times and bubbly episodes, and depending on the different sized bubbles. The second one is on the effects of government spending policies, as the slackness of the economy changes.

Economist is interested in the effects of government policies. However, most literature use linear model to analyze the historical average effect. How about policies have different effects when the economy is varied? In the light of this, the first chapter introduce a novel empirical framework to study generalized state-dependent impulse response functions, a methodology which should have many applications in macroeconomics. Using a flexible semiparametric varying-coefficient model specification, we find the effect of an exogenous tightening monetary shock on stock prices as an increasing function of the size of bubbles. We also employ a two-state threshold local projection method, and find that when the Fed increases the Federal Funds rate, the stock price decreases in normal times but increases during bubbly episodes. The asymmetric response is shown to be driven by both the asymmetric effects on fundamentals and the existence of bubbles, when allowing time-varying risk premium. The state-dependent evidence is more informative in measuring monetary policy effects than linear or time-varying methods, and is also robust to different identification schemes and various definitions of bubbles. This paper points out two important transmission channels of monetary policy on asset price: risk premium and asset bubbles, which are often ignored in theoretical models. On the policy side, our empirical analysis suggests that central banks should be cautious about adopting "leaning against bubble" monetary policies when the bubble size is relatively large.

The conventional interest rate monetary policy analyzed above is ineffective after 2008 due to the zero lower bound constraint. This has brought fiscal policy to the forefront of policy discussions. One key question in the current policy debate is the size of the fiscal multiplier. Following the same methodology as the first work, the second chapter is directly related to this question. We analyze whether the fiscal multiplier differs during different periods, relying on more than one hundred years of U.S. historical data. Using the flexible semiparametric varying coefficient method in the framework of local projections, I particularly model the fiscal multiplier as a function of unemployment rate. The paper shows that the U.S. fiscal multiplier is slightly below one and approximately the same, during periods of slack as compared to normal times. The empirical results suggest that fiscal policy is not necessarily a more powerful tool to stimulate aggregate demand during the Great Recession.

2. THE STATE-DEPENDENT EFFECTS OF MONETARY POLICY ON ASSET BUBBLES

2.1 Introduction

Asset bubbles and financial crises have been recurring phenomena, but what could central banks do to contain bubbles in order to avoid another crisis? Prior to 2008, the Greenspan-Bernanke Federal Reserve followed the "Jackson Hole Consensus," which advocates that central banks should focus on inflation and output gap, while ignoring fluctuations in asset prices.[1] However, because of the severity and duration of the Great Recession, many economists suggest that monetary authority should react to asset bubbles.[2] Yellen (2009) concludes in her speech,

> "[M]onetary policy that leans against bubble expansion may also enhance financial stability by slowing credit booms and lowering overall leverage."

According to the "leaning against bubble" monetary policy, the central bank should target asset prices in order to achieve financial stability, even at the cost of a transitory deviation from the optimal inflation and output gap target. However, Gali (2014) calls into question the theoretical foundations of this policy in the context of an overlapping generations model with nominal rigidities. In particular, the author shows that a systematic increase in interest rates in response to a growing bubble

[1] This pre-crisis consensus is associated with two views which state that asset bubbles are very difficult to identify and measure, and moreover, even if they could be observed, the interest rate would still be too "blunt" an instrument to manage (Gali (2014)). See Bernanke and Gertler (1999) and Bernanke and Gertler (2001) for theoretical foundation and Kohn (2006) for policy recommendation.

[2] Kuttner (2012) offers a detailed empirical assessment to challenge the "Jackson Hole Consensus." He presents two main points: (1) macroeconomic stability does not necessarily guarantee financial stability; (2) financial stability should not be overemphasized because the bursting of an asset bubble can be a disaster to the whole economy.

will actually intensify its fluctuations.[3] Thus, it is still not fully understood how monetary policy should be conducted to react to asset bubbles.

Instead of the normative perspective regarding how monetary authorities should respond to asset bubbles, this paper takes a positive approach in an attempt to shed light on the following interesting questions: Are financial markets reactions to the conventional monetary policy, i.e., changing the Federal Funds Rate (FFR), different during bubbly episodes from what is usually observed in normal times?[4] Do the effects of monetary policy on asset price change as the relative size of the bubble changes? If yes, in what way? As far as we are concerned, in order to make policy recommendations, one should first understand what effect those policies may have on asset prices, and especially their bubble component.

The "Leaning against bubble" principle presumes that the central bank increasing the nominal short-term interest rates will decrease asset prices, which, in effect, reduces the size of an asset bubble. Bernanke and Kuttner (2005) is one of the most influential seminal works to support this view, and they find that a surprise 25-basis-point cut in the Federal Funds rate is associated with about 1% increase in the broad stock index. Gürkaynak, Sack and Swanson (2005) uses the factor model and intraday data to estimate the response of asset prices to factors associated with the Federal Open Market Committee meetings, and find very similar result. However, stock prices rose following the series of FFR increases ending in February 1989, February 1999 and May 2000. Gali and Gambetti (2014) adopt the time-varying coefficient vector-autoregression to rigorously find that the observed effect on stock

[3]The rationale is that a tightening monetary shock will reduce the fundamental price of the asset, but at the same time will increase the bubble component. When the relative size of the bubble is large, the overall effect of increasing the interest rate may drive the asset price up, due to its positive effect on the bubble more than offsetting the negative response of the fundamental component.

[4]Kuttner and Shim (2013) empirically analyze the effect of non-interest rate policies on housing markets using a panel of fifty-seven countries.

prices changes over time, especially during protracted episodes stock prices increase persistently (after a short-run decline) in response to an exogenous contractionary monetary policy shock.[5] They explain that changes in interest rate have a different impact on the two components of the asset price–fundamental and bubble–and the overall effect may change over time as the relative size of the bubble changes. Thus, their finding is essentially the *indirect effect* of bubble sizes on monetary policy outcomes. Moreover, there are other causes and interpretations for this time-varying effect. The effectiveness of monetary policy per se may change over time as the implementation process and communication strategy of monetary authorities evolve.[6] Hence, we aim to *directly* investigate the relationship between monetary policy effects on asset price and bubble sizes.

Our motivation also stems from the lack of economic explanations for the observed equity market response to monetary policy. In fact, changes in policy rate can affect expectations of *future* real interest rates, dividends, risk premium, and bubbles. The first three channels are associated with fundamentals. Bernanke and Kuttner (2005) find the effect of monetary policy on stock price is mainly driven by changes in risk premium, but they rule out the possibility of bubbles. Though Gali and Gambetti (2014) explain that part of the effect is from bubbles, they ignore the important risk premium channel. In this paper, we try to fill this gap by examining all four transmission channels together.

This paper empirically investigates the state-dependent dynamic effects of interest rate policy to capture asset price behavior during normal and bubbly episodes, as well as under different sized bubbles. We start with estimating the local projections

[5]The empirical evidence supports Gali (2014) theory of "*Monetary Policy and Rational Asset Price Bubbles.*

[6]See Primiceri (2005) on time-varying effect of monetary policy on output and Jansen and Zervou (2014) on stock return.

(Jordà (2005), henceforth LP) on U.S. monthly data from 1969: M1 to 2007: M9.[7] As for identification, we use narrative evidence introduced by Romer and Romer (2004) (R&R, hereafter) as exogenous monetary policy shock, to account for the anticipatory behavior of the FOMC.[8][9] Next, we run the two-state threshold local projections as in Owyang, Ramey and Zubairy (2013) to investigate the dynamic effects of monetary policy on stock prices during bubble (B) and normal times (N). We detect and date bubbly episodes as black Monday (1986 M03-1987 M09) and the dot-com bubble (1995 M07- 2001 M08), following Generalized Supreme Augmented Dickey-Fuller (GSADF, in Phillips, Shi and Yu (2013)) test. We also disentangle possible transmission channels of monetary policy on stock price. Last, to generalize our results, we allow the state to be a continuous variable of bubble size instead of using exogenous cutoffs.[10] Specifically, we use semiparametric varying-coefficient model (Jansen, et al. (2008)) in the framework of local projections, to more directly and clearly investigate the state-dependent effect of monetary policy on asset prices as an unspecified smooth function of the relative size of asset bubbles.

In a nutshell, the main results of this paper are as follows: (I) the effect of interest rate policy on asset price is different in two regimes: when the Fed increase FFR the stock price decreases during normal times but increases during bubbly episodes. We

[7]Since this paper focuses on interest rate policies, rather than the broader realm of monetary policy such as macro-prudential instruments, we use data up until the recent financial crisis excluding the periods that zero lower bound constraint the FFR behavior.

[8]The R&R shock is updated by Barakchian and Crowe (2013) and Coibion, et al. (2012). We also use the market based monetary policy shocks by Barakchian and Crowe (2013) to check the robustness of our result.

[9]Our motivation comes from the limitation of traditional identification method in VAR initiated by Christiano, Eichenbaum and Evans (1999), namely "Cholesky decomposition." Both Chen (2007) and Gali and Gambetti (2014) follow that identification scheme and assume that the monetary policy shocks do not contemporaneously affect GDP, dividends, and inflation. While this is the most widely used identification scheme (conceivably due to its easiness in implementation), it suffers from foresight bias (non-fundamentalness problem).

[10]In our benchmark analysis, we use the growth rate of price dividend ratio as the measure of bubble size. For robustness check, we also choose stock volatility as an alternative measure.

allow time-varying risk premium and show that unlike in normal times the effect of monetary policy cannot be fully explained by fundamentals during bubbly episodes, which suggests the existence of bubbles. (II) The effect of conventional monetary policy on asset price changes as the relative size of the bubble varies: the reaction function of stock price to a contractionary monetary shock increases as the bubble component becomes bigger, from negative to positive territory.

Compared to the other linear or time-varying evidence, the state-dependent results may be more relevant and appealing to both theoretical economists and policy makers. From a modeling point of view, this paper supports Gali (2014) theory of "*Monetary Policy and Rational Asset Price Bubbles*," but also points out a missing transmission channel of monetary policy on stock price, which is risk premium. On the policy side, our empirical analysis suggests that central banks should be cautious about adopting "leaning against bubble" monetary policies when the bubble size is relatively large. For example, when the annualized growth rate of price-dividend ratio is more than 100%, the stock price will increase contemporaneously by 1% in response of 100 basis points increase of the Federal Funds rate. Policy makers may need to resort to other unconventional monetary policy or financial policy to contain the asset bubble expansion.

Another contribution of this paper is from the methodology perspective. We propose a novel empirical framework –local projections with semi-parametric varying coefficient approach– to study generalized state-dependent impulse response functions. The model specification is very flexible, data-driven and without imposing any restrictions on the form of impulse response function, which should be of interest to macro-economists and has many applications in other contexts: the effect of monetary policy on output and inflation may change as the overall financial stress varies,[11] fiscal multipliers may be different depending on the slackness of the economy,[12] as well as the changing pattern of the effect of financial shocks as uncertainty changes.

Our empirical results complement previous work on testing the existence of asset price bubble. In a survey paper, Gürkaynak (2005) documents "for almost every paper in the literature that 'finds' a bubble, there is another one that relaxes some assumption on the fundamentals and fits the data equally well without resorting to a bubble." However, this paper does not make any additional assumptions, our model of fundamentals is only based on present value identity. Most importantly, while allowing time-varying risk premium to measure fundamentals, we are still able to show the existence of bubbles.[13]

This paper is also related to the vast literature on the asymmetric effectiveness of monetary policy. Davig (2006), and Jansen and Tsai (2010) examine asymmetries in the impact of monetary policy surprises on stock returns between bull and bear markets. However, they only investigate the immediate effect, rather than the dynamic response of stock prices to unanticipated monetary shocks discovered by our paper. Another set of articles analyze the state-dependent effects of monetary policy on the real economic variables such as output, consumption and investment. Tenreyro and Thwaites (2013) estimate the impulse response of key US macro series to the monetary policy shocks in expansions between in recessions. Dahlhaus (2014) studies the effects of a monetary policy expansion on 108 U.S. macroeconomic and

[11] Dahlhaus (2014) analyzes this phenomenon, but through a smooth-transition factor model.

[12] There is large literature examines the state-dependent fiscal multiplier through nonlinear parametric model (Auerbach and Gorodnichenko (2012), and Owyang, Ramey and Zubairy (2013), among others), and Zhou (2014) investigates the issue following the similar method proposed in this paper.

[13] Note that in the paper, we do not directly calculate fundamental and bubble, alternatively we investigate whether the effect of monetary policy on overall stock price could be fully explained by the effect on its fundamental component. Although Gali and Gambetti (2014) find similar result, the assumption of constant risk premium makes their conclusion less convincing.

financial time series during times of high financial stress versus during normal times. While as stated by Bernanke and Kuttner (2005), the influence of monetary policy instruments on macroeconomic variables is at best indirect, this paper should help us understand the state-dependent policy transmission mechanism.

2.2 Theoretical Issues

There is growing anecdotal evidence regarding bubbles, but what is an asset price bubble from an economist's point of view? Do asset bubbles really exist in financial markets? Generally speaking, if the price of an asset is more than its "fundamental", because investors expect to be able to sell the asset at an even higher price in the future, then there is an bubble (Gürkaynak (2005), Malliaris (2012)).[14]

2.2.1 Present Value Model

Following Cochrane (2005), we start by introducing the Present-Value identity:

$$1 \equiv R_{t+1}^{-1} R_{t+1} \equiv R_{t+1}^{-1} \frac{P_{t+1} + D_{t+1}}{P_t}, \tag{2.1}$$

where R is the gross simple return, P_t is the after-dividend price of the asset, and d_t is the payoff (dividend) received from the asset. In the context of the stock market, P_t is the stock, and d_t is dividend; while for the real estate market, P_t and d_t are house price and rent price respectively.[15]Hence,

$$\frac{P_t}{D_t} = R_{t+1}^{-1}(1 + \frac{P_{t+1}}{D_{t+1}}) \frac{D_{t+1}}{D_t}. \tag{2.2}$$

[14]Throughout this paper, we only talk about rational bubbles, which should be positive by its definition. Another strand of literature relying on behavior models, can generate "irrational bubbles," which can be neither positive or negative. Reader interested in this area are referred to Vissing-Jorgensen (2003) and Scherbina (2013) for surveys.

[15]Alternatively, P_t maybe price of a mine, and d_t is the value of ore unearthed every period.

Taking conditional expectation,

$$P_t = E_t \frac{(P_{t+1} + D_{t+1})}{R_{t+1}}, \tag{2.3}$$

and iterating forward by repeatedly substitute out future prices and using the Law of Iterated Expectations, we have,

$$P_t = E_t[\sum_{k=1}^{\infty}(\prod_{j=0}^{k-1}(1/R_{t+j}))D_{t+k}] + \lim_{k\to\infty} E_t[\prod_{j=0}^{k-1}(1/R_{t+j})P_{t+k}]. \tag{2.4}$$

The first term on the right-hand side of the equation is the fundamental component of asset prices, defined as the present discounted value of future dividends, can be written as:

$$P_t^F \equiv E_t \sum_{k=1}^{\infty}(\prod_{j=0}^{k-1}(1/R_{t+j}))D_{t+k}. \tag{2.5}$$

Moreover, the second term is the expected discounted value of the future (infinity) stock price, and we can define is as the "bubble" part Q^B of asset price.

$$P_t^B \equiv \lim_{k\to\infty} E_t[\prod_{j=0}^{k-1}(1/R_{t+j})P_{t+k}]. \tag{2.6}$$

Note that the traditional transversality condition rules out the existence of bubble, which relies on the assumption that the expected discounted future stock price converges to zero as the horizon goes to infinity. However, the theory of rational asset pricing bubble relax the above assumption. The rationale is that investors appear to be betting that other investors would drive prices even higher in the future. The bubble part is assumed to follow the below process, which is derived from the ratio between current and subsequent period's bubble:

$$P_t^B = E_t \frac{P_{t+1}^B}{R_t}. \tag{2.7}$$

This is the key equation for "rational bubble" theory, and the adjective "rational" represents the bubble is entirely consistent with rational expectations.

The most important feature of this Present-Value model (PVM) is that there is not any assumptions required: all is based on accounting identity. Readers may find this advantage easily when we compare the PVM with the partial equilibrium model discussed in the next subsection.

2.2.2 Partial Equilibrium Model

This section briefly reviews the standard asset pricing model augmented with bubble component, which is a partial equilibrium version of the model discussed in Gali (2014) as well as Gali and Gambetti (2014).

As in Gürkaynak (2005), we maximize the expected utility consumption, $u(c)$, in an endowment economy,

$$\max \quad E_t \sum_{i=1}^{\infty} \beta^i u(C_{t+i}) \tag{2.8}$$

$$s.t.\ C_{t+i} = Y_{t+i} + (P_{t+i} + D_{t+i})X_{t+i} - P_{t+i}X_{t+i+1}, \tag{2.9}$$

where Y_t is the endowment, β is the discount factor, X_t is the storable asset, P is the asset price (after dividend), and D is the dividend.

The first order condition (FOC) of the optimization problem is:

$$E_t \beta u'(C_{t+i})[P_{t+i} + D_{t+i}] = E_t \beta u'(C_{t+i-1}) P_{t+i-1}. \tag{2.10}$$

One of the basic assumption in the traditional asset pricing model follows,

Assumption 1 *The utility function is CRRA (constant relative risk aversion) form.*

It is assumed implicitly or explicitly that utility is linear, which implies constant marginal utility and risk neutrality.

Assumption 2 *There is a risk-less bond available in zero net supply with one period net interest rate, r^f, which is also called risk free rate.*

Two assumptions together with no-arbitrage condition imply the FOC can be derived as:

$$E_t(P_{t+i-1}) = \frac{1}{1+r^f} E_t(P_{t+i} + D_{t+i}). \tag{2.11}$$

This first-degree difference equation can be iterated forward to obtain its solution. Basically, asset price P is the sum of "fundamental" component P^F and "bubble" part P^B, that is:

$$P_t = P_t^F + P_t^B. \tag{2.12}$$

The fundamental component, defined as the present discounted value of future dividends, can be written as:

$$P_t^F \equiv E_t \sum_{k=1}^{\infty} \left(\prod_{j=0}^{k-1} \frac{1}{1+r_{t+j}^f} \right) D_{t+k}. \tag{2.13}$$

The bubble part is assumed to follow:

$$P_t^B = E_t \frac{P_{t+1}^B}{1+r_t^f}. \tag{2.14}$$

Note that,

$$r_t = r_t^f + rp_t, \tag{2.15}$$

where r_t is the net discount rate, r^f is the risk-free rate, and rp_t is risk premium. The partial equilibrium model is essentially the restricted version of PVM, in the sense that the risk premium is always equal to zero.

2.2.3 Impulse Response Function

What is the effect of monetary policy on asset price? The answer to this question is much more complicated than it seems to be. The "Leaning against bubble" principle that is consistent with the conventional wisdom, assumes that the central bank increasing the nominal short-term interest rates will decrease the asset prices. However, is it supported by the theory?

After log-linearization, and simple derivation (see Appendix), we can get the predicted response of the fundamental component:

$$\frac{\partial p^F_{t+k}}{\partial \varepsilon^m_t} = \sum_{j=o}^{\infty} \Lambda^j [(1-\Lambda)\frac{\partial d_{t+k+j+1}}{\partial \varepsilon^m_t} - \frac{\partial r^f_{t+k+j}}{\partial \varepsilon^m_t} - \frac{\partial rp_{t+k+j+1}}{\partial \varepsilon^m_t}], \tag{2.16}$$

where ε^m_t is monetary policy shock at period t and $k = 0, 1, 2, ...$ denotes the period after the initial shock. $\Lambda \equiv \Gamma/R$, with Λ and Γ denote gross growth rate of dividend and interest along a balanced growth path, respectively.

Through liquidity effect, an exogenous tightening monetary policy will cause a rise in the real interest rate, i.e. $\partial r_{t+k}/\partial \varepsilon^m_t > 0$. Moreover, as discussed in Gali and Gambetti (2014), both conventional wisdom and economic theory point out that contractionary monetary policies decrease the dividends, that is $\partial d_{t+k}/\partial \varepsilon^m_t \leq 0$. However, the effect of monetary policy on risk premium is not straightforward. If the effect on risk premium is positive or slightly negative, the effect on fundamentals may be negative; but if the effect on risk premium is largely and persistently negative, the impulse response on fundamentals could be positive. Thus, we are not certain even about the sign of monetary policy effects on fundamentals.

However, the effect of monetary policy on the bubble part of asset prices is not straightforward due to the lack of theoretical foundation. Under the "leaning against wind" monetary policy, a contractionary monetary policy should cause a decline in the size of the bubble, i.e. we expect:

$$\frac{\partial P_{t+k}^{B}}{\partial \varepsilon_t^m} \leq 0. \tag{2.17}$$

On the contrary, the theory of rational asset price bubbles in Gali (2014) has a different prediction. In a rational expectations equilibrium, both $P_t R_t = E_t[D_{t+1} + P_{t+1}]$ and $P_t^F R_t = E_t[D_{t+1} + P_{t+1}^F]$ must hold, so the bubble part must satisfy:

$$P_t^B R_t = E_t P_{t+1}^B. \tag{2.18}$$

Hence the expected growth rate of the bubble component will increase in response to a rise in real interest rate. After several steps of derivation, Gali and Gambetti (2014) have shown the dynamic effect of monetary policy on the bubble component is given by:

$$\frac{\partial p_{t+k}^{B}}{\partial \varepsilon_t^m} = \begin{cases} \psi_t \frac{\partial r_t}{\partial \varepsilon_t^m} & \text{if } k = 0, \\ \psi_t \frac{\partial r_t}{\partial \varepsilon_t^m} + \sum_{j=0}^{k-1} \frac{\partial r_{t+j}}{\partial \varepsilon_t^m} & \text{if } k > 0, \end{cases} \tag{2.19}$$

where ψ_t is a possible random parameter, but the existing economics theory can hardly pin down its sign or size. Thus the dynamic impulse response of bubble part to such shock is indeterminate.

To sum up, the effect of monetary policy on asset price is ambiguous, at least from the above theoretical model. This motivates us to use econometric tools to examine the issue empirically.

2.3 Econometric Methodology for Two-State Analysis

I will first introduce the identification scheme to monetary policy, and then the econometric model used to examine the two-state impulse response function, as well as the statistical test to date multiple bubbles.

2.3.1 R&R's Measure of Monetary Policy Shocks

How to identify monetary policy shock is always an important question for macroeconomists. After Sims (1980), using structural VAR to calculate the shocks is the most widely used approach. However, this method relies on several strong assumptions. The first one is the data generating process in the real world should follow VAR, which is somehow unrealistic. The other is that in order to derive the structural shocks from reduced innovation, one needs to impose additional assumptions, including short-run restrictions, long-run restrictions, and sign-restrictions. Romer and Romer (2004) develops a new narrative measure of U.S. monetary policy shocks for the period of 1969 to 1996 which is relatively free of endogenous and anticipatory movements. More specifically, they estimate a Taylor type of rule with the Greenbook forecast as control variables, and extract residuals s_t to be referred to as the exogenous monetary policy shock:

$$\begin{aligned}\Delta f_t = \alpha + \beta f b_m + \sum_{i=-1}^{2} \gamma_i F_t \Delta y_{mi} + \sum_{i=-1}^{2} \lambda_i (F_t \Delta y_{mi} - F_t \Delta y_{m-1,i}) + \\ \sum_{i=-1}^{2} \varphi_i F_t \Delta \pi_{mi} + \sum_{i=-1}^{2} \theta_i (F_t \Delta \pi_{mi} - F_t \Delta \pi_{m-1,i}) + \rho F_t u_{m0} + s_t,\end{aligned} \tag{2.20}$$

where Δf_t is the intended change in the FFR decided upon at the FOMC meeting at time t, and fb_m is the level of the intended funds rate before any changes associated with meeting m. Other control variables include $F_t \Delta y$, $F_t \Delta \pi$ and $F_t u$, representing the Greenbook forecasts of real output growth, inflation and the unemployment rate. Following Romer and Romer (2004), the i subscript refers to the horizon of the forecast, where -1 and 0 are the past and the current quarter, and 1 and 2 are one and two quarters ahead, respectively.

The R&R shock measure is expected to account for anticipation effect bias and free of model misspecification. While I am not the first to turn to R&R series, this paper is the first to use the narrative evidence to examine the response of asset prices to monetary policy shocks.

2.3.2 Linear Local Projections

Jordà (2005) introduces a new way to do estimation and inference of impulse responses, namely local projection (LP). Compared to VAR, LP is simpler in implementation and more robust to misspecification, thus it has become better received by the economics world recently.[16] For its application, please see Auerbach and Gorodnichenko (2012) and Owyang, Ramey and Zubairy (2013) on fiscal policy, Tenreyro and Thwaites (2013) on monetary policy, Chong, Jordà and Taylor (2012) on exchange rate economics, and Jordà, Schularick and Taylor (2013) on credit business cycle. Let y_t, p_t^i, p_t^c, i_t, p_t, and d_t represent industrial production (IP), CPI, the commodity price index, the Federal Funds Rate (FFR), the S&P 500 stock index (real), and its dividend (real) respectively, then I further define $Y_t \equiv [\Delta y_t, \Delta d_t, \Delta p_t^i, \Delta p_t^c, i_t, \Delta p_t]'$.[17] I start by running a set of regressions for each

[16] There is no restriction on the shape of the impulse response function, which makes the result more data-driven. Since in reality, the data generating process is not necessary be VAR. Moreover, we do not need to estimate a vector of equations when using LP, but just the variable of interest as dependent variable.

[17] All the variables are converted into log term, and then taken the first difference before going into the regressions.

horizon h as follows:

$$\Delta p_{t+h} = \alpha_h + \psi_h(L)Y_{t-1} + \gamma_h s_t + \epsilon_{t+h}, \tag{2.21}$$

where h is from 0 to 12. Since in this model s_t is our monetary policy shock measure, γ_h is the estimated impulse response coefficients that of interest:

$$IR(s,h) = \gamma_h. \tag{2.22}$$

As for the serial correlation problem, we use the Newey-West correction for the standard errors. Let $\hat{\Sigma}$ be the estimated heteroskedasticity and autocorrelation (HAC), variance-covariance matrix of the coefficients γ; then a 95-percent confidence interval for the impulse response can be constructed approximately as $1.95 \pm \hat{\Sigma}$.

However, the potential drawbacks for LP is that the estimates are sometimes erratic, due to the fact of the loss of efficiency. This is mainly because that we do not impose a strong constraint on the shape of impulse response function.[18] Moreover, at longer horizon there is sometimes oscillations emerged in the impulse responses.[19] Fortunately, this problem is not quite a concern for us, since we are only interested in the short-run behavior of the effect of monetary policy on asset price.

2.3.3 Two-State Threshold Local Projections

Another important advantage of LP is that it easily accommodates with highly nonlinear and flexible specifications. Following Owyang, Ramey and Zubairy (2013) (henceforth ORZ), I adopt two-state threshold LP technique to investigate whether the effects of monetary policy on asset prices are different between normal times and

[18]Note that for the case of employing VAR to calculate impulse response function, the moving average (MA) representation that people mainly rely on is highly restricted.

[19]See Ramey and Zubairy (2014) for a detailed comparison between LP and standard VAR.

the bubbly episodes.[20]Similar to ORZ, I estimate a sequence of regressions for each horizon h using threshold dummy variables in the context of LP:

$$\Delta p_{t+h} = I_{t-1}[\alpha_{Nh} + \psi_{Nh}(L)Y_{t-1} + \gamma_{Nh}s_t] + (1 - I_{t-1})[\alpha_{Bh} + \psi_{Bh}(L)Y_{t-1} + \gamma_{Bh}s_t] + \epsilon_{t+h}, \tag{2.23}$$

where L denotes polynomials in the lag operator. I is a dummy variable which takes the value of one when the economy is in the normal times (N), and zero when it is in bubble states (B). We allow all of the coefficients to vary according to the state of the economy. The mechanism to separate N and B will be introduced below. Through the lens of two-state threshold LP method, we are able to examine the dynamic responses of stock prices to monetary policy shocks in two different periods, and capture any potential variations between the two. We are able to formally test whether the effect of monetary policy is statistically different during the two periods.[21]

2.3.4 Testing for Multiple Bubbles

Even though asset price bubbles are generally rather difficult to detect and measure, recent advances in econometric detection mechanisms have shown success in identifying and dating financial bubbles. Phillips, Wu and Yu (2011) is one of the recent papers contributing to this issue. In this project, I will follow Phillips, Shi and Yu (2013)'s method, which is a new recursive testing procedure that can effectively locate the dates of multiple bubble events. The basic idea of Generalized Supreme Augmented Dickey-Fuller (GSADF) test is a rolling window right-sided ADF unit root test with a double-sup window selection criterion, in an attempt to find explosive

[20]Owyang, Ramey and Zubairy (2013) use two-state local projections method to calculate fiscal multipliers during periods of slack versus in the normal times.

[21]On the contrary, most of the results in Gali and Gambetti (2014) are statistically insignificant, which may render their conclusions less persuasive.

behaviors of the asset price while taking into account the fundamental value.[22]The reduced-form approach to detect bubbles can be written as:

$$\Delta z_t = \alpha_{r_1,r_2} + (\rho_{r_1,r_2} - 1) z_{t-1} + \sum_{i=1}^{k} \psi^i_{r_1,r_2} \Delta z_{t-i} + \varepsilon_t, \tag{2.24}$$

where z is the price-dividend ratio for the S&P 500 index, k is the lag order and ε_t is i.i.d. with 0 mean and $\sigma^2_{r_1,r_2}$ variance. In particular, this is a rolling window regression where the subsample begins at the r_1^{th} fraction of the total sample (T) and ends at the r_2^{th} fraction.[23]

The unit root null hypothesis is a random walk process:

$$H_0 : \rho_{r_1,r_2} = 1, \tag{2.25}$$

and the right-tailed alternative hypothesis is the explosive behavior:

$$H_1 : \rho_{r_1,r_2} > 1. \tag{2.26}$$

The interested reader may refer to Phillips, Shi and Yu (2013) for a clear description regarding their rolling-window test and data-stamping strategies.

After we run the forward recursive regressions then implement the GSADF test, we are able to divide the whole sample periods into two states: normal (N) and bubble (B) times.

[22]As discussed in Phillips, Shi and Yu (2013) and Yiu, Yu and Jin (2013), the proposed bubble stamping method is a sufficient condition for identifying bubbles.

[23]r_ω is the fractional window size, and $r_2 = r_1 + r_\omega$.

2.4 Data Description

We use monthly data in this paper. When analyzing financial markets, a quarter is too long a time because stock returns on a daily, or even intra-day basis, and agents often react quickly to public policy news. Moreover, monetary policy in U.S. is conducted eight times a year, which is more frequently than a quarter.

The World Bank Commodity Price Index (Non-energy) comes from the World Bank Dataset, and the S&P 500 dividend is from Robert J. Shiller's website.[24]All the other variables featured in the LP, industrial production (IP), CPI, FFR, and S&P 500 index are obtained from the Federal Reserve Economic Data (FRED) at the Federal Reserve at Saint Louis. The narrative evidence of monetary policy shock is originally constructed by Romer and Romer (2004), and I am able to obtain the updated version from Barakchian and Crowe (2013) and Coibion, et al. (2012) data set. Since this paper focuses on interest rate policies, rather than the broader realm of monetary policy such as macro-prudential instruments, our data sample is restricted from 1969 to 2008.[25]

2.5 Two-State Evidence

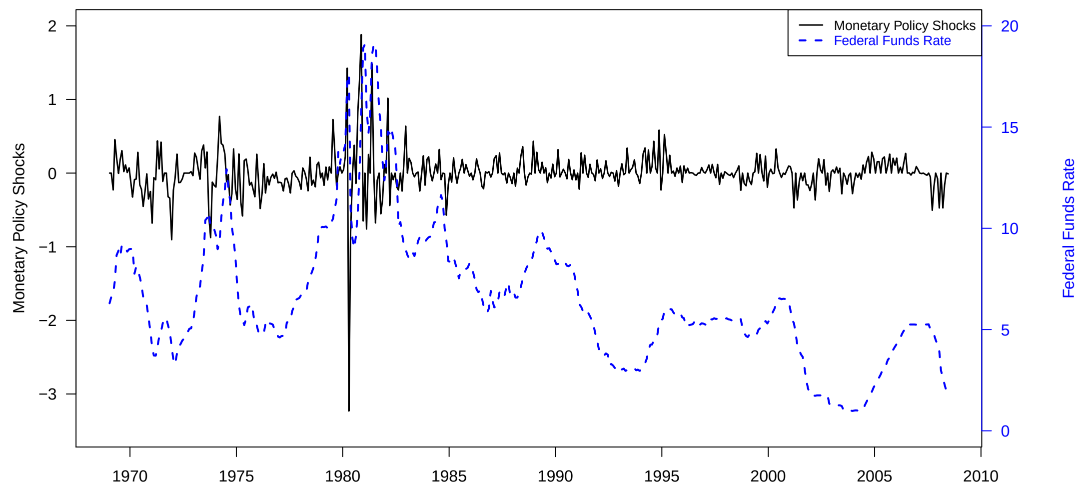
In this section, we report the main empirical evidence obtained from the two-state model.

2.5.1 Measure of Monetary Policy Shocks

We use the updated version of R&R monetary policy shocks by Barakchian and Crowe (2013), which is from January 1969 to June 2008, along with the FFR. One can see from Figure 2.1 that the volatility of the shocks change dramatically from pre- to post-1983 which suggests an adjustment in Federal Reserve behaviors as the

[24]http://www.econ.yale.edu/ shiller/data.htm
[25]The zero lower bound on the FFR since the crisis rendered data after 2008 uninformative.

Figure 2.1: R&R Measure of Monetary Policy Shock and the Federal Funds Rate



Volcker era came. The most fluctuated period is 1979 to 1982 when the Federal Reserve moved to a non-borrowed reserves operating procedure. In the cases when the Federal Reserve explicitly targeted the FFR – between 1974 to September 1979 and the entire period after the mid-1980's – the R&R monetary policy measure is relatively stable.

2.5.2 Bubbly Episodes

Figure 2.2 is the time series graph for the S&P 500 stock index, and its dividend, while Figure 2.3 is the GSADF test to detect multiple bubbles. A visual inspection of the real monthly S&P 500 index and its dividends already suggest that the bigger volatility in stock prices can hardly be explained by fundamentals. Especially, there is accelerating upswing in stock price starting from 1995, but we do not see such pattern in the dividend data. In Figure 2.3, the blue line is the backward SADF sequence, and the red dashed line is the 95% critical value sequence. The shadowed

Figure 2.2: S&P 500 Index and Dividends (Normalized), from 1968 to 2008

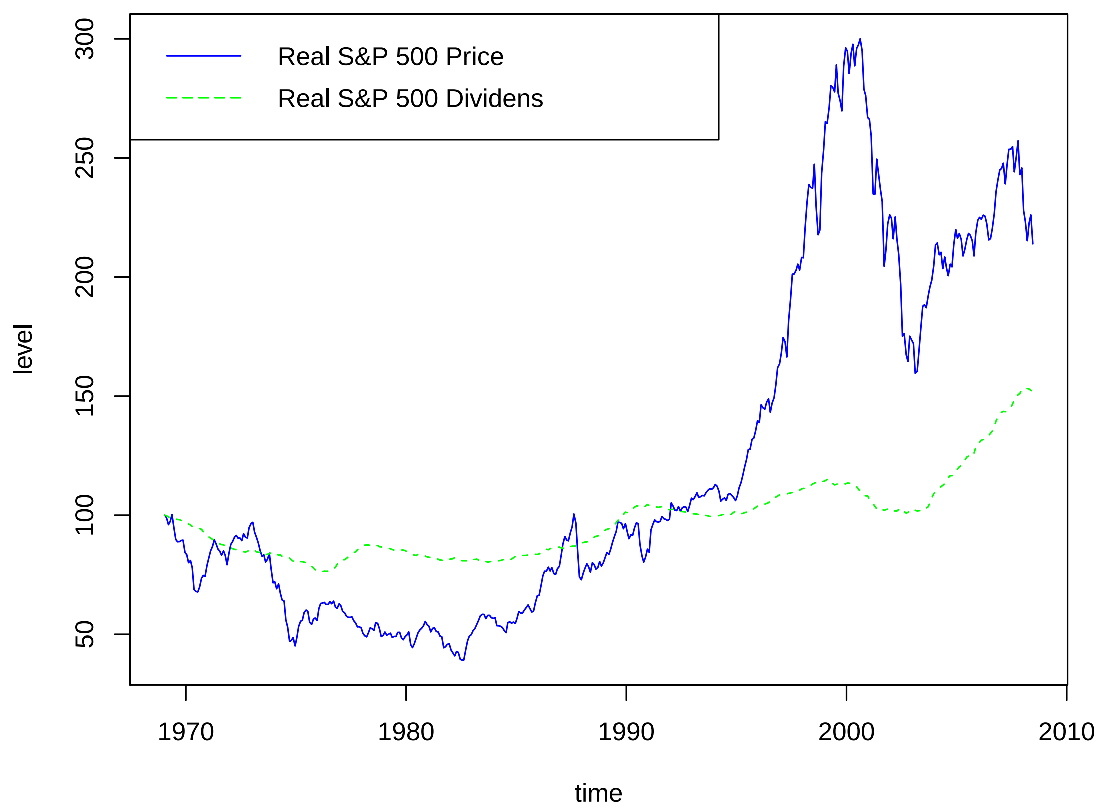
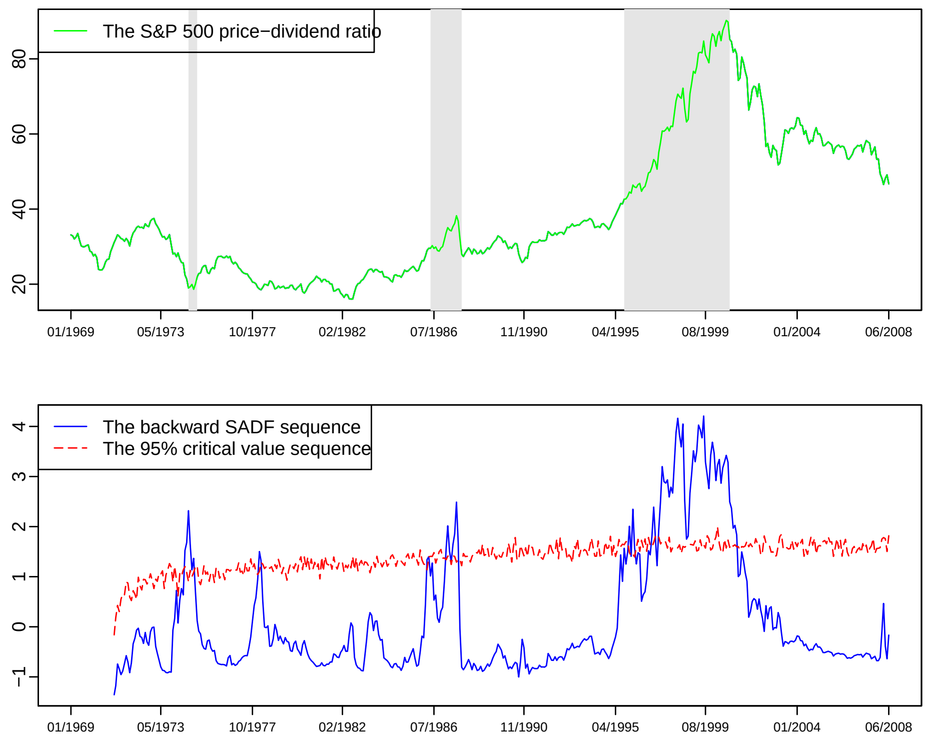


Figure 2.3: Date-Stamping Multiple Bubbles in the S&P 500 Price-Dividend Ratio: the GSADF Test



area represents the bubbly episodes where the GSADF statistic is above its 95 percent confidence interval. In summary, the identified bubble periods are: black Monday in October 1987 (1986 M03-1987 M09), the dot-com bubble (1995 M07-2001 M08) and the 1974 stock market crash (1974 M07-M12, which is consistent with the well-known historical episodes for bubbles.)[26] Our dataset is nearly forty years, and if we sum up the bubbly episodes together there are around eight years which is 20% of the whole sample period.

2.5.3 Linear and Two-State Evidence

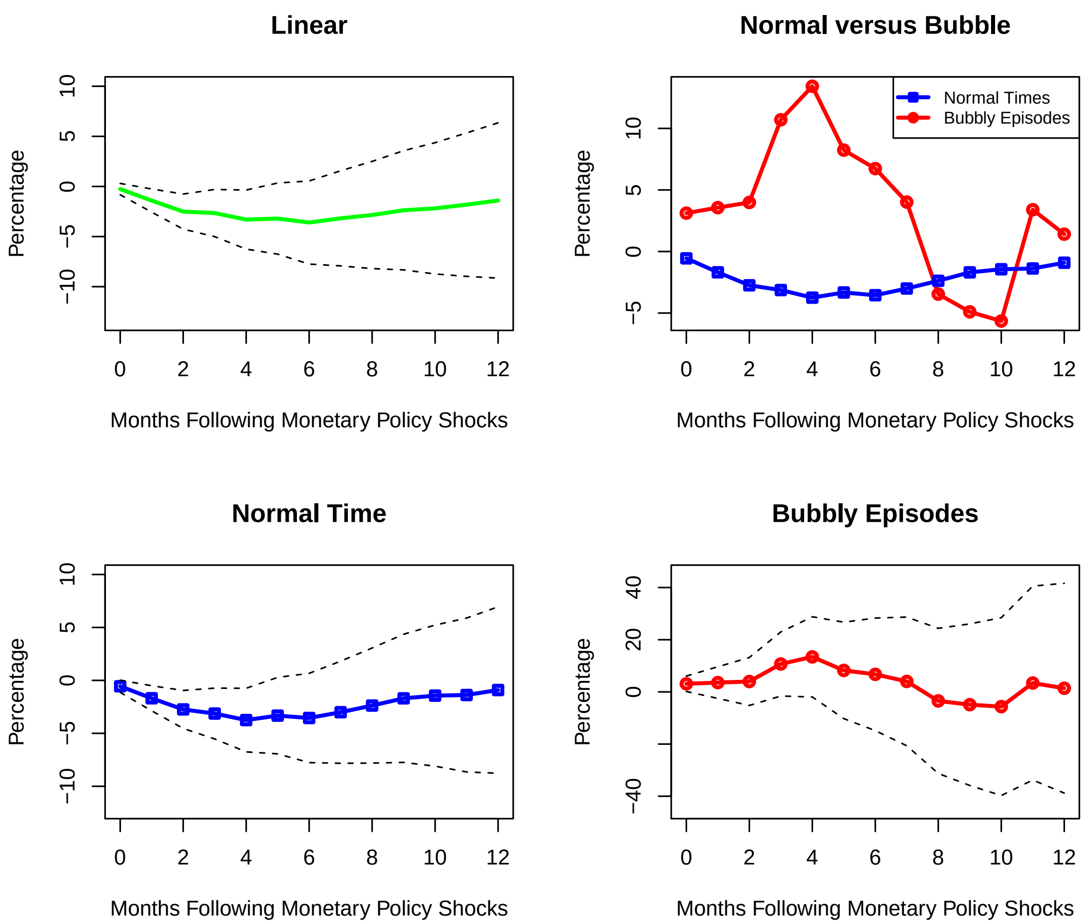
In Figure 2.4, we show the impulse response of S&P 500 index to a monetary policy shock by linear, and two-state LP models. In the linear model, the S&P 500 declines initially, then reaches its maximum in 6 months, but recovers in about one year. This finding is consistent with the result obtained from constant parameter VAR in Gali and Gambetti (2014), but our evidence is statistically significant, at least for the first few months.

It is apparent that the graphs representing the state-dependent effect between the two different periods - N and B - are vastly different. The responses in normal times are not very different from those of the whole sample period estimated by the linear model, but with a larger initial decline after a positive monetary policy shock. However, the responses of the S&P 500 to monetary policy shocks during bubble phases are striking and very different from its counterpart in normal times: the price increases for about 5 months (rather than decrease), and then recovers and returns to a positive territory.[27] The estimated impulse response function for the bubble phases is similar to 1984Q4-1987Q3 (years before the Black Monday stock market crash) as

[26]Some people may argue that the market crash period is not really a bubbly episode, but including it or not does not affect our results much.

[27]However, the confidence bands are too large to reject the no-effect hypothesis.

Figure 2.4: S&P 500 Response to Exogenous Monetary Policy Shocks, Linear and Two-State Local Projection



well as 1997Q1-1999Q4 (just before the dot com bubble burst), both of which are studied in Gali and Gambetti (2014).[28] This result may also call the "leaning against bubble" policies into question.

2.6 Two-State Interpretation

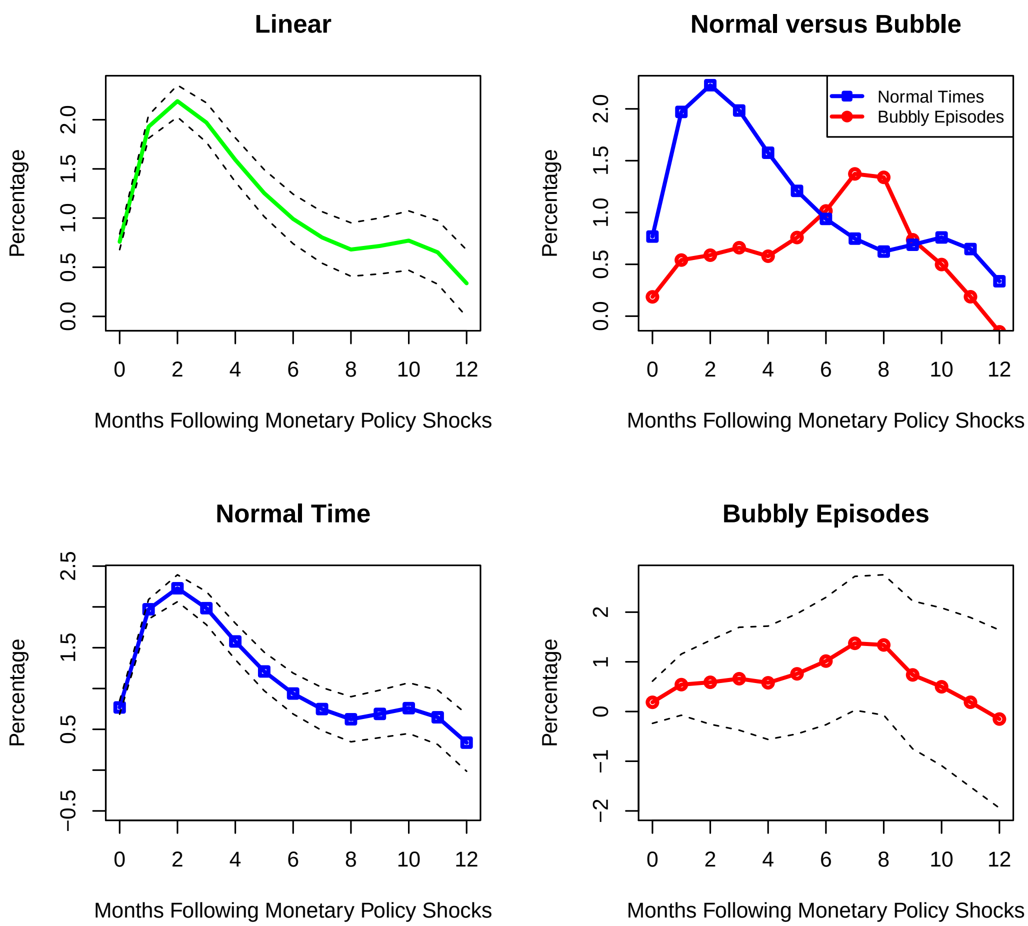
How to explain the stock market's reaction to monetary policies, especially the observed state-dependent evidence? Are they associated with a change in expected future dividends, real interest rates, equity premiums, or through some other channels, such as bubbles? How does monetary policy affect the bubble component of the stock price? In this section we turn to these difficult questions, and disentangle the various channels through a more structured approach.

2.6.1 Federal Funds Rate

First of all, the systematic component of monetary policy itself might behave differently. One natural question arises: are the dynamic effects of monetary shocks on the Federal Funds rate different during bubbly episodes from those in normal times? It is interesting and important to understand the policy implementation in the two different regimes. Figure 2.5 shows the impulse response of the Federal Funds rate to monetary contraction, during bubbly episodes and during normal times, as well as through the linear model. We do not find substantial variation of dynamic responses between the two states, which suggests that the central bank does not implement different monetary policies in two states.

[28] In their paper, the authors do not explicitly define which periods are bubbly phases, instead they rely on conventional wisdom of the most famous crises-black Monday and Dot.com crisis-and refer samples before the crises as bubble times.

Figure 2.5: The Federal Funds Rate Response to Exogenous Monetary Policy Shocks, Linear and Two-State Threshold Local Projection



2.6.2 Fundamental Component

We generalize the impulse response function of the fundamental component to exogenous monetary shocks, by introducing the role of risk premium.[29] A growing body of literature shows that risk premium is very important in explaining the effect of monetary policy on financial markets. Bernanke and Kuttner (2005) empirically find that the effects of unanticipated monetary policy actions on expected excess returns account for the largest part of the response of stock price, but only a small part of effects is directly from changes in the risk-free rate.[30] From a theoretical side, Drechsler, Savov and Schnabl (2014) build a dynamic heterogeneous-agent asset pricing model to investigate the transmission channel between monetary policy and risk premium. However, Gali and Gambetti (2014) do not provide substantial quantitative results on the effect of monetary policy on excess return. Figure 2.6 is the impulse response of excess return to monetary contraction, by linear and two-state LP methods. Similar to the effect on stock price, there is substantial variation in dynamic responses between the two states. We will use this calculated response function to further derive the impulse response for the fundamental component later. We have also calculated the two-state impulse response function of monetary policy on real interest rate and dividend, respectively. It is not statistically different between normal times and bubbly episodes from Figure 2.7 to Figure 2.8: the real interest rate increases, while the dividend decreases in response of a monetary tightening shock.

Next, we present the results of the impulse response function for the "subjective" fundamental component, which allows time varying risk premium. From the second

[29]Gali and Gambetti (2014) mention this issue in their "Alternative Interpretation" section, but without any supporting empirical evidence.

[30]Thought the paper, we use risk premium and expected excess return interchangeably.

Figure 2.6: Excess Return Response to Exogenous Monetary Policy Shocks, Linear and Two-State Threshold Local Projection

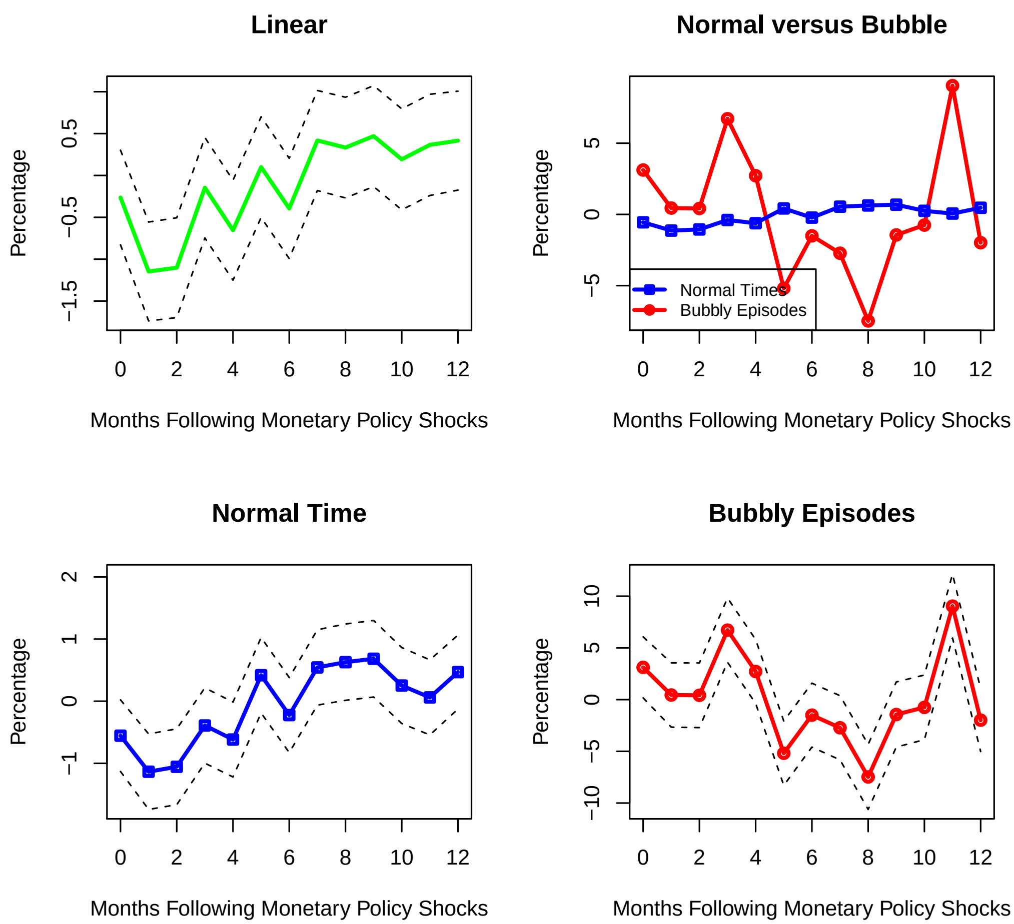


Figure 2.7: Real Interest Rate Response to Exogenous Monetary Policy Shocks, Linear and Two-State Threshold Local Projection

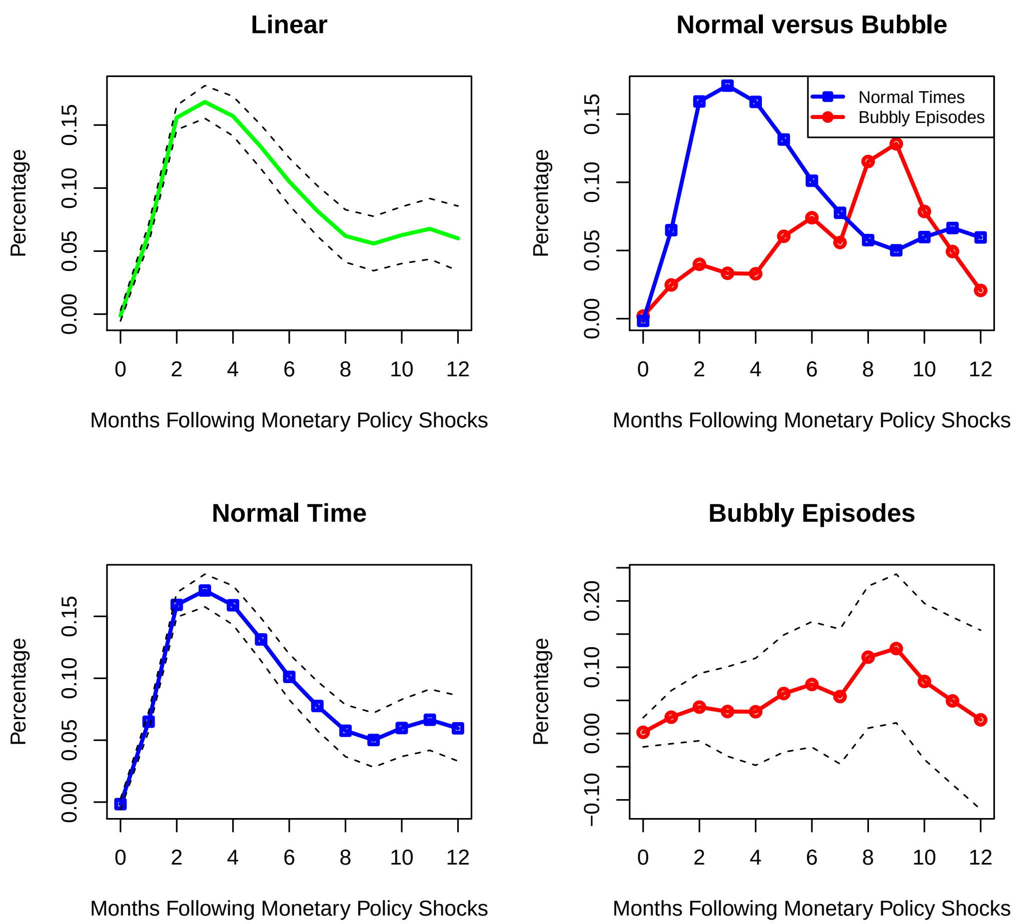


Figure 2.8: Dividend Response to Exogenous Monetary Policy Shocks, Linear and Two-State Threshold Local Projection

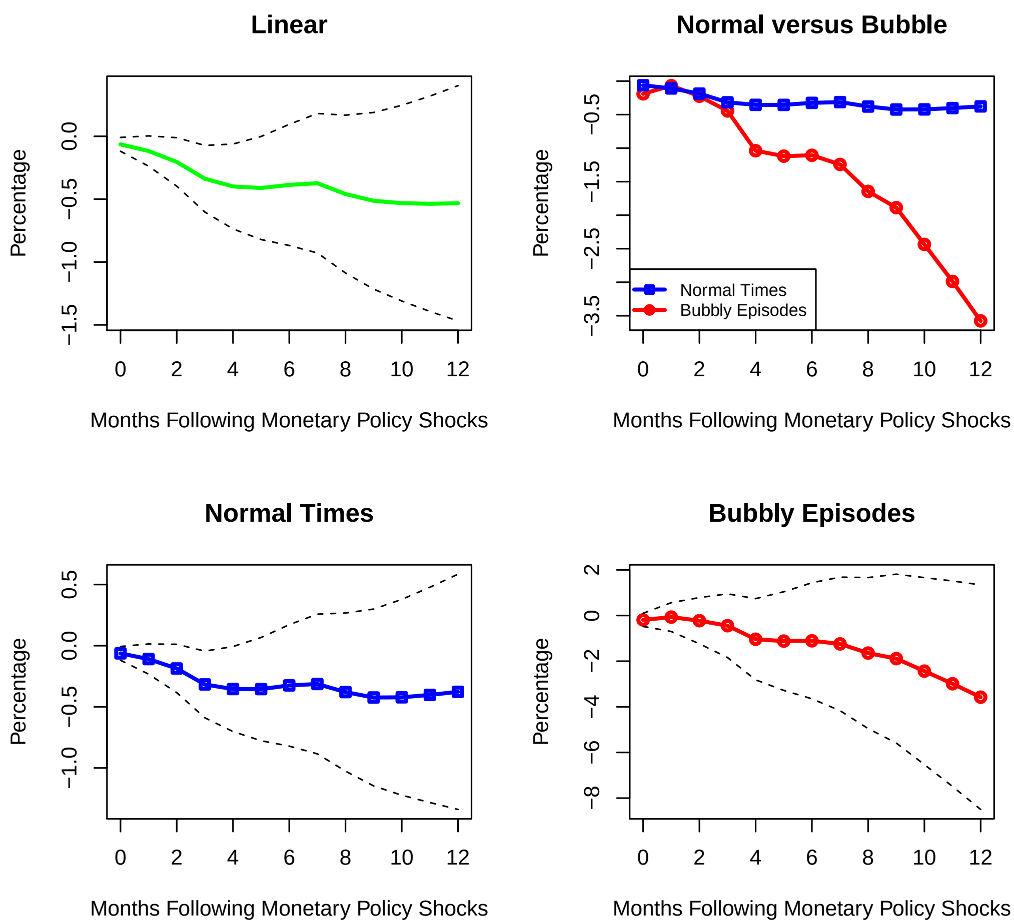
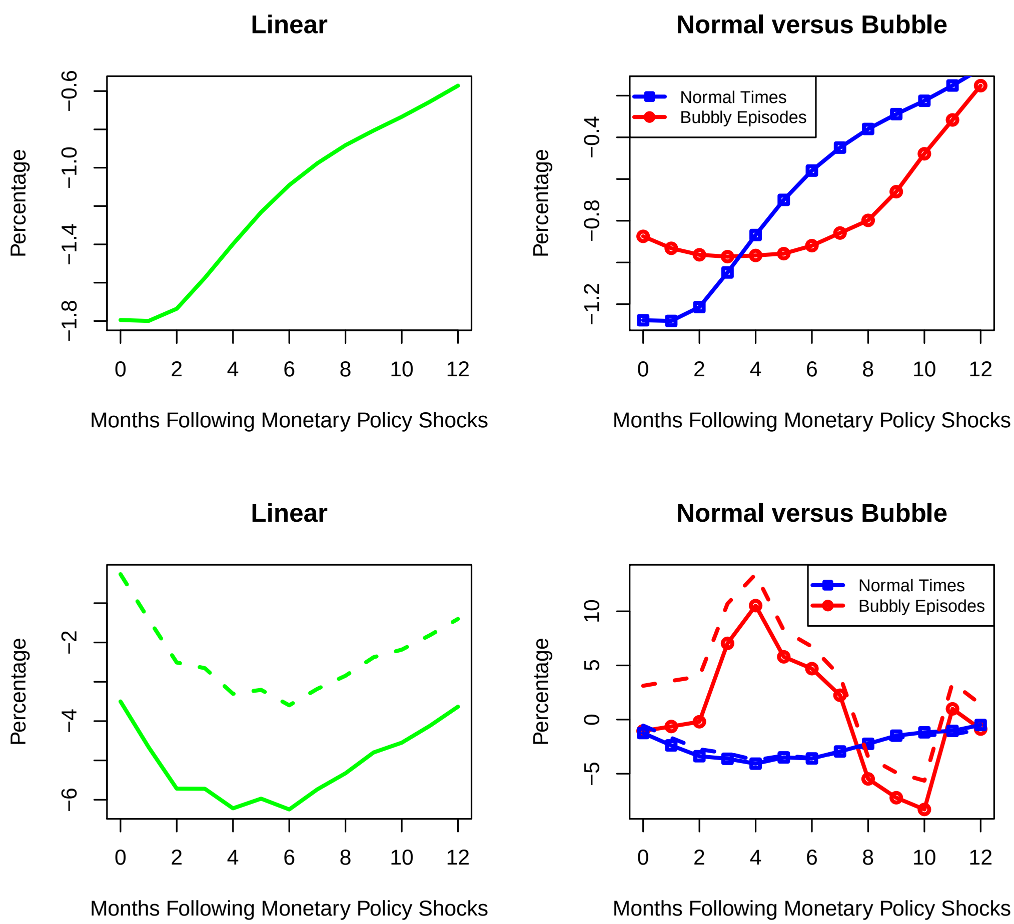


Figure 2.9: Fundamental Response to Exogenous Monetary Policy Shocks, Linear and Two-State Threshold Local Projection



row of Figure 2.9, it is obvious to see that the dynamic effect of monetary policy on the fundamental component differs in bubbly episodes and in normal times. This result contradicts Gali and Gambetti (2014), where they find the impulse response for the fundamentals is largely unchanged over time. However, their finding is based on the constant risk premium assumption. We also calculate the result using their method:

$$\frac{\partial p_{t+k}^{F_r}}{\partial \varepsilon_t^m} = \sum_{j=o}^{\infty} \Lambda^j [(1-\Lambda)\frac{\partial d_{t+k+j+1}}{\partial \varepsilon_t^m} - \frac{\partial r_{t+k+j}^f}{\partial \varepsilon_t^m}]. \tag{2.27}$$

We call it the "objective" fundamental. If only relying on the first row of Figure 2.9, one may get the misleading finding that there is no state-dependent effect on the fundamental component, which is similar to Gali and Gambetti (2014)'s result. This bias is mainly driven by ignoring the role of time-varying risk premium when explaining the effect of monetary policy.

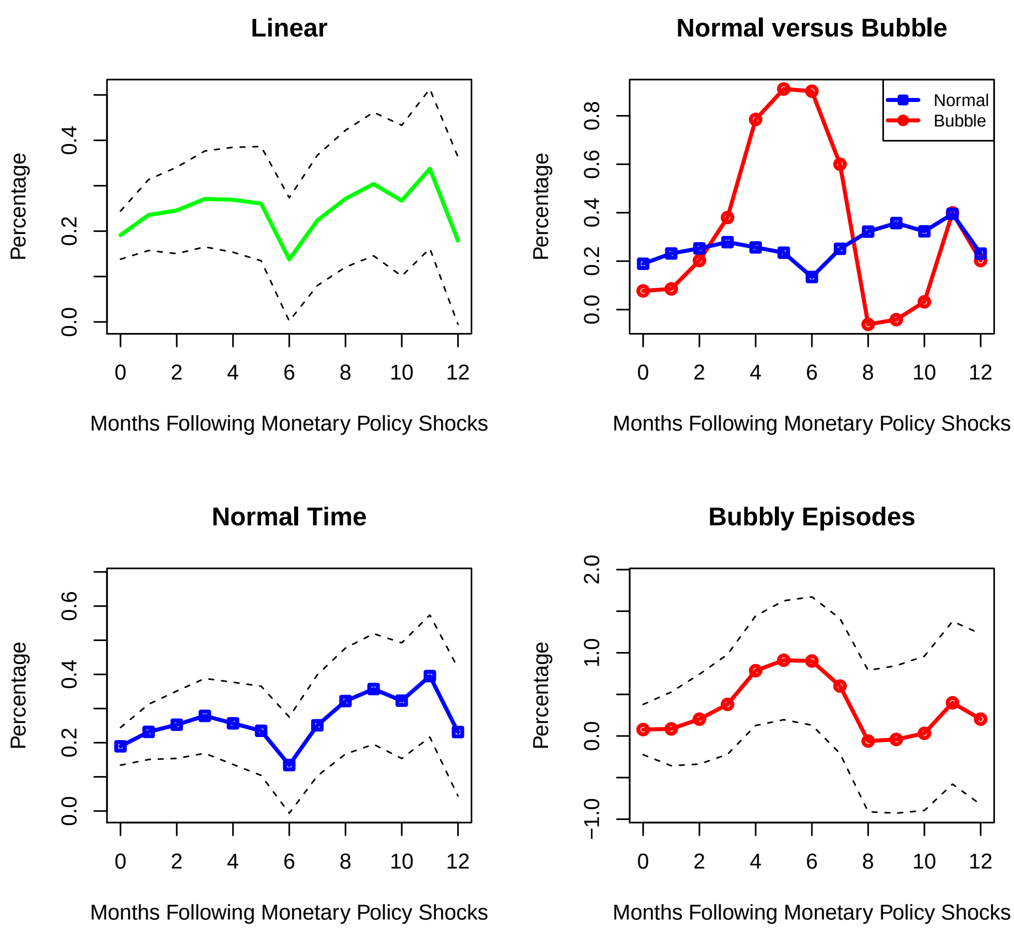
Moreover, comparing the effect of monetary policy shock on the overall stock price with the "subjective" fundamentals, we still find that there exists a positive gap between them, but only during the bubbly episodes. The interpretation is that the effect of monetary contraction is positive on the bubbles, and the bubble component is negligible during normal times but relatively large during bubbly episodes. This finding can be explained by Gali (2014)'s "monetary policy and rational asset price bubbles" theory.

In summary, the state-dependent effect we have observed comes not only from the existence of bubble component, but also from the assymetric effects of monetary policy on fundamentals.

2.6.3 Long Term Interest Rate

Motivated by Gali and Gambetti (2014), we analyze the response of ten-year US treasury bond to exogenous monetary policy shocks using our state-dependent method. Another hypothesis could also explain the observed evidence: If people mistakenly anticipate that short term rates will remain low for a sufficiently long period of time, the monetary tightening may coexisted with a simultaneous decline in the long term interest rate. Similar to Gali and Gambetti (2014), we find that

Figure 2.10: Long-Term Rate Response to Exogenous Monetary Policy Shocks, Linear and Two-State Threshold Local Projection



in response of a contractionary monetary policy shock, the long-term interest rate rises persistently both during bubbly episodes and normal times. Thus, the above hypothesis is not supported by our empirical result.

2.7 Generalized State-Dependent Analysis

In this section, we first introduce a novel empirical framework to analyze generalized state-dependent impulse response function, and then use this proposed method to directly capture the effect of monetary policy on stock prices depending on bubble sizes.

2.7.1 Generalized State-Dependent Impulse Response Function: Semiparametric Varying-Coefficient Model with LP

The semiparametric varying coefficient model, also known as the functional-coefficient regression model, is becoming more widely used in economics nowadays.[31] As Li and Racine (2007) stated, its advantage is allowing more flexibility in functional forms than either a linear model or many parametric nonlinear models, and at the same time avoiding much of the "curse of dimensionality" problem that occurs in fully nonparametric analysis. Jansen, et al. (2008) is one of the papers using semiparametric specification to analyze the macroeconomics issue: in particular, they examine the role of fiscal policy plays in US asset markets. Li Lin and Hsiao (2014) use similar methods on international economics, trying to test Purchasing Power Parity hypothesis in a more flexible way.

Jordà (2005) points out that the functional form of the control variables in LP may include any parametric, semiparametric and non-parametric approximation. Ten years later, little work, if any, has been conducted relating LP with semiparametric or non-parametric forms. This paper is the first to use semiparametric varying coefficient model in the framework of LP to evaluate the state-dependent impulse

[31] One may refer to Cai Fan and Yao (2000) and Li et al. (2002) for descriptions of estimation methods and asymptotic distribution in detail.

response function. More specifically, I run a set of semiparametric varying coefficient regressions for each horizon h as follows:

$$\Delta p_{t+h} = \alpha_h(z_t) + \psi_h(z_t)(L)Y_{t-1} + \gamma_h(z_t)s_t + u_{ht}, \tag{2.28}$$

where z is the state dependent variable and we choose it as the measure of bubble size in this paper. In contrast, α, ψ and γ are unspecified smooth functions of z, rather than the estimated average parameters in parametric models. Instead of defining each state as either normal or bubble using exogenous cutoffs, I allow it to be a data-driven continuous variable (price dividend ratio) that describes the size of the bubble. This way the state dependent effects should be captured through state-dependent function if they indeed exist. In this paper, the function form of interest is $\gamma(z_t)$, which represents the effect of monetary policy shocks on asset prices as a function of price dividend ratio.[32]

Under the assumption that model is correctly specified, we must have:

$$E(u_t|Y_{t-1}, s_t, z_t) = 0. \tag{2.29}$$

For simplicity, we can rewrite the model as:

$$\Delta p_{t+h} = X_t'\beta(z_t) + u_{ht}, \tag{2.30}$$

where $X_t \equiv (ones, Y_{t-1}, s_t)$, and $\beta(z)$ is a vector of coefficient functions. Pre-multiplying both sides of the model by X_t, then taking conditional expectation,

[32]A higher growth rate of price dividend ratio z is likely to be correlated with a bigger asset bubble.

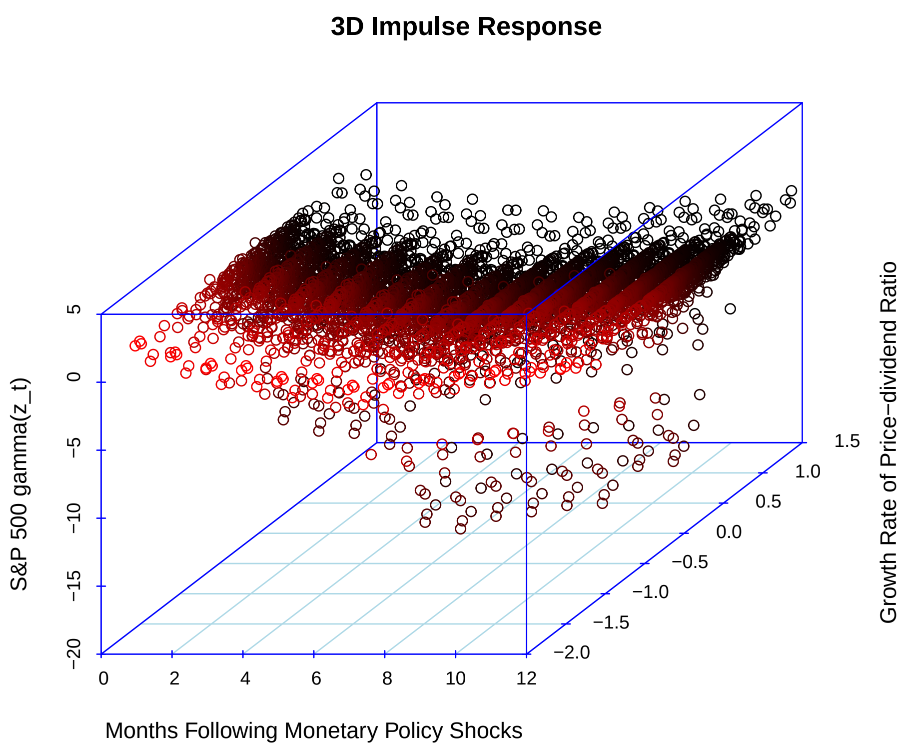
we can get:

$$\beta(z) = [E(X_t * X_t'|z_t = z)]^{-1} E(X_t * X_t'|z_t = z). \tag{2.31}$$

One can use either local-constant or local-linear kernel estimation (see Appendix) to obtain a feasible estimator of $\beta(z)$. Interested reader may refer to Li and Racine (2007) for a comprehensive and thorough introduction of semiparametric and nonparametric estimation.

2.7.2 Generalized State-Dependent Evidence

Figure 2.11: S&P 500 Response to Exogenous Monetary Policy Shocks, Semiparametric Varying-Coefficient Local Projection, Three-Dimensional.



We present the impulse response function of SVC-LP, which is a function of time as well as the growth rate of price dividend ratio, in a three dimensional Figure 2.11. Compared to the traditional impulse response function calculated by parametric models (either linear or nonlinear), what we report here is different in the sense that the response is not only changing over time, but also a function of a continuous state variable. We can see a substantial curvature in the surface of this figure, which suggests the state-dependent evidence. For a clearer picture about how the effect changes as the size of bubble varies, we also report the reaction function for selected periods in Figure 2.12: contemporaneous, four, eight and twelve months after. All the other period by period figures from the contemporaneous month to twelve months after are in the Appendix B). Generally speaking, the state-dependent effect of monetary policy on stock prices can be captured by an increasing function of the growth rate of price dividend ratio: the estimated effect rises as the price-dividend ratio increases, at least until twelve months after the initial shock. Figure 2.13 illustrate the changing patterns of stock price response by showing the average impulse responses over four different regimes: $-1 < \Delta z < -0.5$; $-0.5 < \Delta z < 0$; $0 < \Delta z < 0.5$; $0.5 < \Delta z < 1$. In this graph, from bottom to top, as the bubble size increases, the negative effect of a monetary tightening on stock price become smaller.

2.8 Robustness Checks

In this section, we implement several robustness checks: alternative definition of bubble, and different identification methods for monetary policy shocks, to compare our results with other literature.

2.8.1 Alternative Measure of Bubble

Suggested by Basco and Crespo (2014), the volatility of the stock price is an alternative proxy of the bubble component. The rationale is that volatility is larger

Figure 2.12: S&P 500 Response to Exogenous Monetary Policy Shocks, Semiparametric Varying-Coefficient Local Projection, Selected Periods.

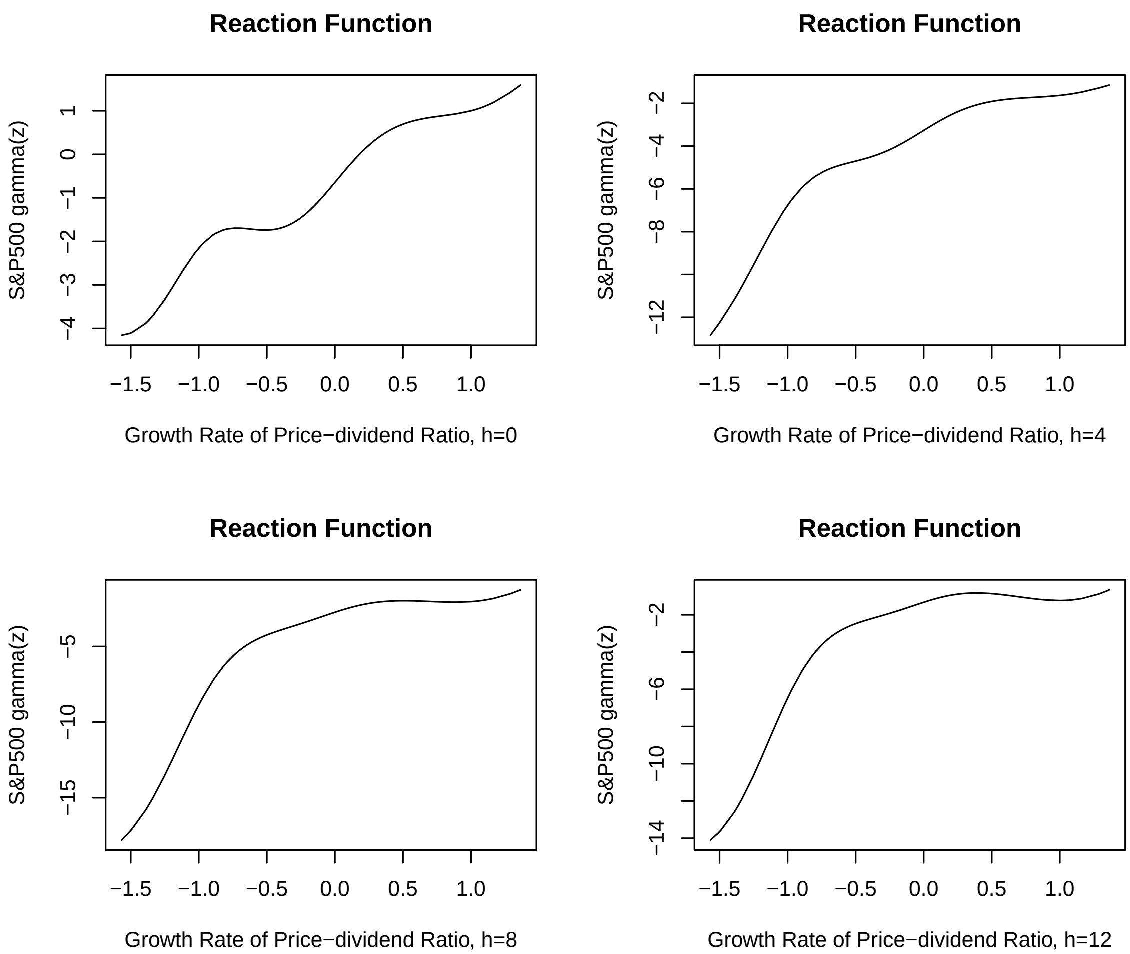
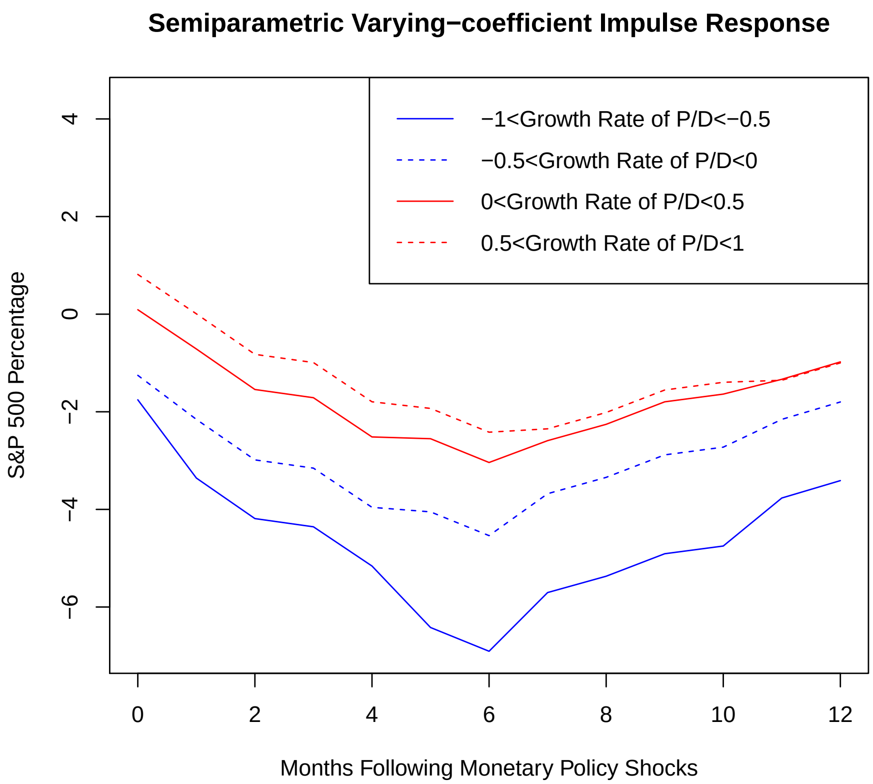
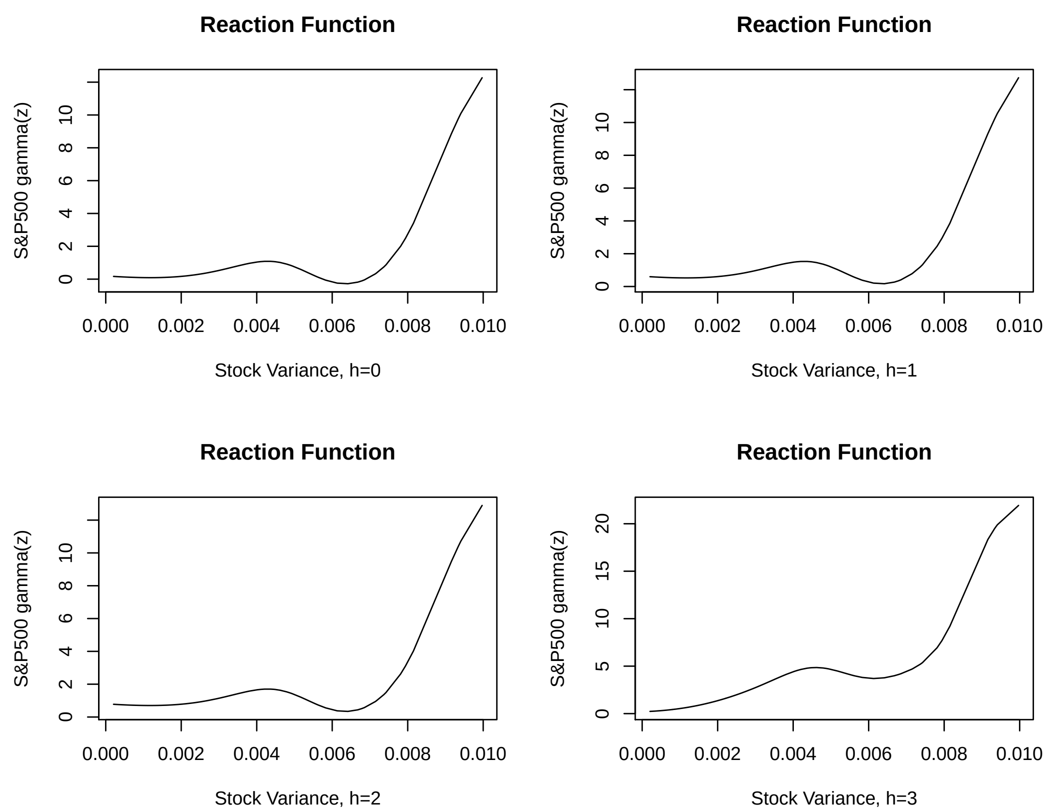


Figure 2.13: S&P 500 Response to Exogenous Monetary Policy Shocks, Semiparametric Varying-Coefficient Local Projection, Different Ranges of the Growth Rate of Price Dividend Ratio.



in stock prices with a bubble than without one, due to the fact that the fundamental component is less volatile than the bubble part. We use the sum of squared daily returns on the S&P 500 to capture stock variance, which is widely accepted in finance literature (Welch and Goyal (2008)).[33] Please refer to appendix for the time series graph of the realized volatility of S&P 500. Figure 2.14 shows the reaction function

Figure 2.14: S&P 500 Response to Exogenous Monetary Policy Shocks, Semiparametric Varying-Coefficient Local Projection, Alternative Bubble Measure, Selected Periods.



of stock price to an exogenous tightening monetary shock as a function of stock

[33]It is also called realized variance or realized volatility.

variance, from the present to three months after the initial shock. We can see a apparent increasing function, which is similar to the benchmark result.

2.8.2 Market-Based Monetary Policy Shock

We utilize the market based monetary policy shocks by Barakchian and Crowe (2013) to check the robustness of our results, and we refer to their new shock as B&C shocks. Similar to Kuttner (2001), they extract information from Federal Funds futures to measure unexpected changes in interest rate policy, but through a factor model. The intuition of this identification scheme is that movements in Fed Funds futures contract prices on the days of FOMC announcement should capture the "surprise" component of monetary policy actions. We can see the B&C shock series from 1988 to 2008 in Appendix A.[34] The state-dependent results we obtained in Figure 2.15 are largely unchanged, which strengthens our previous findings.

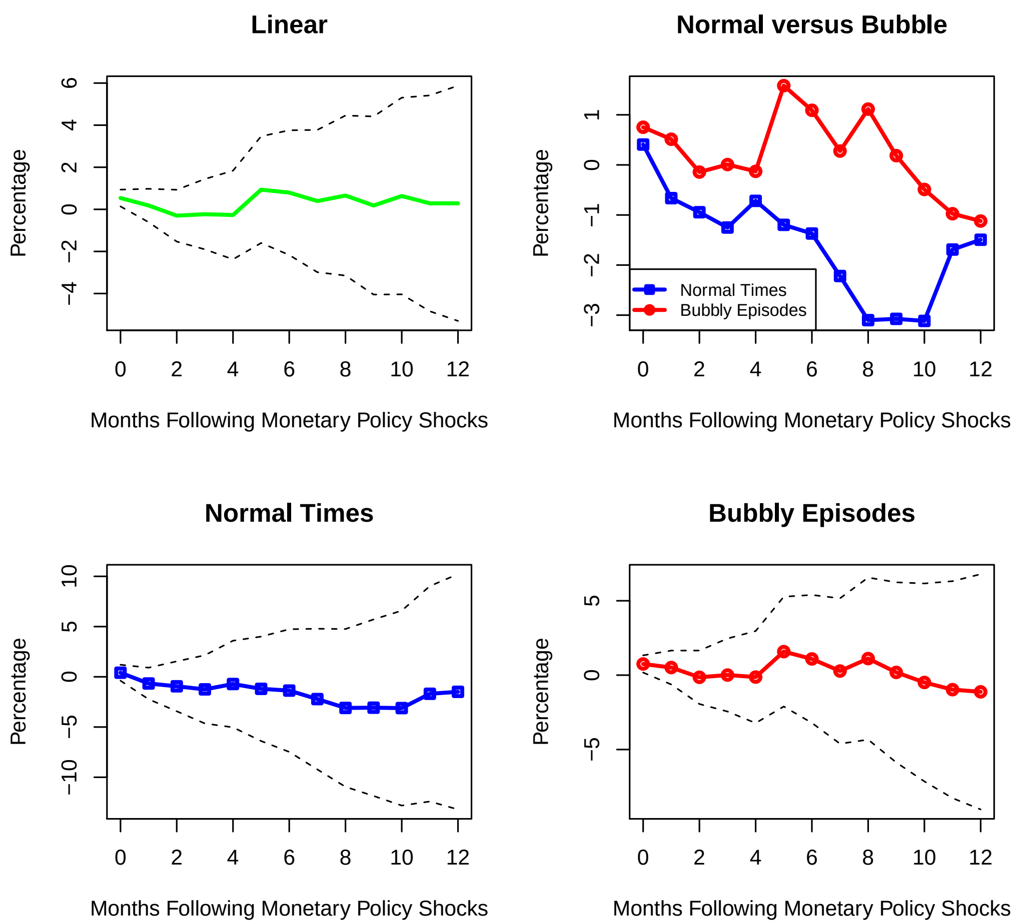
2.9 Conclusion

The "Great Recession" has brought both asset bubbles and monetary policy to the forefront of policy discussions. Should central banks follow the "Jackson Hole Consensus" as the pre-2008 periods, or should they adopt the "leaning against bubble" monetary policy to aggressively contain asset bubbles? Through a positive analysis, this paper shall shed new light on this heated debate.

To this point, we have provided substantial empirical evidence on the response of asset prices to conventional interest rate policy from a state-dependent perspective. The paper can be viewed as an effort to enhance our understanding of the link between monetary policy and financial markets. Particularly, we evaluate the empirical merits of "*leaning against bubble*" principle which believes that stock prices

[34] The sample restricted to post 1988 is mainly because the Fed Funds futures market only started trading in October 1988.

Figure 2.15: S&P 500 Response to Exogenous Monetary Policy (B&C) Shocks, Linear and Two-States Threshold Local Projection



would decrease after short term nominal interest rate increases. In short, our main results, which are shown to be robust to different definitions of bubble and various identification schemes of monetary policy shocks, are listed as follows:

1. Variations do exist in state-dependent effect between normal times and bubbly episodes. Using two-state local projection method, we find the dynamic responses of stock prices to exogenous contractionary monetary shocks are different: when the Fed increases the Federal Funds Rate, the stock price will decrease during normal times but increase during bubble phases. We allow time-varying risk premium and show that this result is driven by both the asymmetric effect on fundamentals and the existence of bubbles.

2. Generalized State-dependent effect is based on the size of the bubble component. Through semiparametric varying-coefficient model in the framework of local projection, we are able to see a clearer picture of the state dependent effect of the interest rate policy on asset prices. This effect, as we identified, takes the shape of an increasing function of the size of bubbles, the range of which goes from negative to positive.

Our findings support the theory of Gali (2014) "*Monetary Policy and Asset Price bubbles*", in the sense that the degree of reactions of asset prices to unanticipated interest rate changes may differ as the relative bubble size varies. However, we also point out another important channel to explain the effect of monetary policy on asset price that is often ignored in theoretical models, which is risk premium. From a policy perspective, our state-dependent evidence suggests that central banks should be cautious about "leaning against bubble" monetary policy, especially when the bubble size is relatively large. Another contribution of the paper is that we have proposed an empirical framework to study generalized state-dependent impulse response functions, a methodology which should have many applications in other contexts of macroeconomics.

However, there are several questions that are not fully answered by our paper. Through which channels do the monetary policies affect asset price bubbles? What is the optimal monetary policy to contain asset bubbles? To answer these questions, a theoretical model with financial frictions and imperfect information would be required in order to understand the transmission mechanism behind our empirical evidence. Thus, future exploration of the structural link between monetary policy and asset markets should be high on the research agenda.

3. IS THE FISCAL MULTIPLIER STATE-DEPENDENT? A SEMIPARAMETRIC ANALYSIS.

3.1 Introduction

The conventional interest rate monetary policy is ineffective after 2008 due to the zero lower bound constraint. This has brought fiscal policy to the forefront of policy discussions. One key question in the current policy debate is the size of the fiscal multiplier. Following a novel methodology, this paper is directly related to this question. I analyze whether the fiscal multiplier differs during different periods, relying on more than one hundred years of U.S. historical data.

Is fiscal policy a good tool to stimulate the economy when in recession? The answer to this question is far from straightforward, since there is no consensus about the size of the fiscal multiplier. One extreme case is that Robert Barro (2009 Wall Street Journal op-ed) argues that fiscal multipliers are close to zero, while Christina Romer estimates multipliers as high as 1.6 by the $787 billion stimulus package approved by the U.S. Congress in February 2009. Different economic models have opposite statements on the effect of fiscal policy. On the one side, the standard neoclassical model suggests that increasing in government spending will crowd out both consumption and investment. On the other side, the traditional Keynesian model predicts that consumption and investment should respond positively to positive spending shocks. Most interestingly, if government spending shocks affect output through Keynesian channels, we expect larger expansionary effects when the economy has significant resource slack than when it is operating at or near full capacity. Intuitively, when the economy has slack, there is more room for the expansionary fiscal policy to stimulate the economy, than to crowd out private consumption and investment. The purpose of this paper, is to directly test the interesting and fundamental issue.

The continuous state analysis is absent in the literature, we therefore propose a novel empirical framework –local projections with semi-parametric varying coefficient approach– to study generalized state-dependent impulse response functions. The model specification is very flexible, data-driven and without imposing any restrictions on the form of impulse response function, which should be of interest to macro-economists and has many applications in other contexts: the effect of monetary policy on output and inflation may change as the overall financial stress varies, [1] as well as the changing pattern of the effect of financial shocks as uncertainty changes.

This paper contributes to the empirical literature on investigating whether government spending multipliers differ according to the different state of the economy. Using the flexible semi-parametric varying coefficient method in the framework of local projections, I directly model the fiscal multiplier as a function of various state variables. The paper shows that the U.S. fiscal multiplier is slightly below one and approximately the same, during periods of slack, recession or zero-lower bound as compared to normal times. The empirical results suggest that fiscal policy is not necessarily a more powerful tool to stimulate aggregate demand during the Great Recession.

3.2 Econometric Methodology

I will introduce the econometric models used in this paper below.

3.2.1 Linear Local Projections

Jordà (2005) introduces a new way to do estimation and inference of impulse responses, namely local projection (LP). Compared to VAR, LP is simpler in implementation and more robust to misspecification, thus it has become better received

[1] Dahlhaus (2014) analyzes this phenomenon, but through a smooth-transition factor model.

by the economics world recently.[2]For its application, please see Auerbach and Gorodnichenko (2012) and Owyang, Ramey and Zubairy (2013) on fiscal policy, Tenreyro and Thwaites (2013) on monetary policy, Chong, Jordà and Taylor (2012) on exchange rate economics, and Jordà, Schularick and Taylor (2013) on credit business cycle. Let y_t, and g_t represent the GDP, and government spending respectively, then I further define $Y_t \equiv [\Delta y_t, \Delta g_t]'$.[3]I start by running a set of regressions for each horizon h as follows:

$$x_{t+h} = \alpha_h + \psi_h(L)Y_{t-1} + \gamma_h s_t + \epsilon_{t+h}, \tag{3.1}$$

where h is from 0 to 12. x is the variable of interest. Since in this model s_t is our government spending news shock, γ_h is the estimated impulse response coefficients:

$$IR(s,h) = \gamma_h. \tag{3.2}$$

As for the serial correlation problem, we use the Newey-West correction for the standard errors. Let $\hat{\Sigma}$ be the estimated heteroskedasticity and autocorrelation (HAC), variance-covariance matrix of the coefficients γ; then a 95-percent confidence interval for the impulse response can be constructed approximately as $1.95 \pm \hat{\Sigma}$.

However, the potential drawbacks for LP is that the estimates are sometimes erratic, due to the fact of the loss of efficiency. This is mainly because that we do not impose a strong constraint on the shape of impulse response function.[4] Moreover,

[2]There is no restriction on the shape of the impulse response function, which makes the result more data-driven. Since in reality, the data generating process is not necessary be VAR. Moreover, we do not need to estimate a vector of equations when using LP, but just the variable of interest as dependent variable.

[3]All the variables are converted into log term, and then taken the first difference before going into the regressions.

[4]Note that for the case of employing VAR to calculate impulse response function, the moving average (MA) representation that people mainly rely on is highly restricted.

at longer horizon there is sometimes oscillations emerged in the impulse responses.[5] Fortunately, this problem is not quite a concern for us, since we are only interested in the short-run behavior of the effect of monetary policy on asset price.

3.2.2 Two-State Threshold Local Projections

Another important advantage of LP is that it easily accommodates with highly nonlinear and flexible specifications. Following Owyang, Ramey and Zubairy (2013) (henceforth ORZ), I adopt two-state threshold LP technique to investigate whether the effects of monetary policy on asset prices are different between normal times and the bubbly episodes.[6] Similar to ORZ, I estimate a sequence of regressions for each horizon h using threshold dummy variables in the context of LP:

$$\Delta x_{t+h} = I_{t-1}[\alpha_{Nh} + \psi_{Nh}(L)Y_{t-1} + \gamma_{Nh}s_t] + (1 - I_{t-1})[\alpha_{Bh} + \psi_{Bh}(L)Y_{t-1} + \gamma_{Bh}s_t] + \epsilon_{t+h}, \tag{3.3}$$

where L denotes polynomials in the lag operator. I is a dummy variable which can capture the different state of the economy.

3.2.3 Generalized State-Dependent Impulse Response Function: Semiparametric Varying-Coefficient Model with LP

In this subsection, we introduce a new methodology to estimate the state-dependent impulse response functions, which is the generalization of the previous linear or non-linear models. The semiparametric varying coefficient model, also known as the functional-coefficient regression model, is becoming more widely used in economics nowadays.[7] As Li and Racine (2007) stated, its advantage is allowing more flexibility

[5] See Ramey and Zubairy (2014) for a detailed comparison between LP and standard VAR.

[6] Owyang, Ramey and Zubairy (2013) use two-state local projections method to calculate fiscal multipliers during periods of slack versus in the normal times.

[7] One may refer to Cai Fan and Yao (2000) and Li et al. (2002) for descriptions of estimation methods and asymptotic distribution in detail.

in functional forms than either a linear model or many parametric nonlinear models, and at the same time avoiding much of the "curse of dimensionality" problem that occurs in fully nonparametric analysis. Jansen, et al. (2008) is one of the papers using semiparametric specification to analyze the macroeconomics issue: in particular, they examine the role of fiscal policy plays in US asset markets. Li Lin and Hsiao (2014) use similar methods on international economics, trying to test Purchasing Power Parity hypothesis in a more flexible way.

Jordà (2005) points out that the functional form of the control variables in LP may include any parametric, semiparametric and non-parametric approximation. Ten years later, little work, if any, has been conducted relating LP with semiparametric or non-parametric forms. This paper is the first to use semiparametric varying coefficient model in the framework of LP to evaluate the state-dependent impulse response function. More specifically, I run a set of semiparametric varying coefficient regressions for each horizon h as follows:

$$x_{t+h} = \alpha_h(z_t) + \psi_h(z_t)(L)Y_{t-1} + \gamma_h(z_t)s_t + u_{ht}, \tag{3.4}$$

where z is the state dependent variable and we choose it as the measure of slackness in this paper. In contrast, α, ψ and γ are unspecified smooth functions of z, rather than the estimated average parameters in parametric models. Instead of defining each state as either normal or bubble using exogenous cutoffs, I allow it to be a data-driven continuous variable (unemployment rate) that describes the slackness of the economy. This way the state dependent effects should be captured through state-dependent function if they indeed exist. In this paper, the function form of interest is $\gamma(z_t)$, which represents the effect of fiscal policy shocks on the real GDP as a function of unemployment rate.[8]

Under the assumption that model is correctly specified, we must have:

$$E(u_t|Y_{t-1}, s_t, z_t) = 0 \tag{3.5}$$

For simplicity, we can rewrite the model as:

$$x_{t+h} = X_t'\beta(z_t) + u_{ht}, \tag{3.6}$$

where $X_t \equiv (ones, Y_{t-1}, s_t)$, and $\beta(z)$ is a vector of coefficient functions. Premultiplying both sides of the model by X_t, then taking conditional expectation, we can get:

$$\beta(z) = [E(X_t * X_t'|z_t = z)]^{-1}E(X_t * X_t'|z_t = z). \tag{3.7}$$

One can use either local-constant or local-linear kernel estimation (see Appendix) to obtain a feasible estimator of $\beta(z)$. Interested reader may refer to Li and Racine (2007) for a comprehensive and thorough introduction of semiparametric and nonparametric estimation.

3.2.4 Measure of Fiscal Multiplier

The government spending fiscal multiplier (FM) is the ratio of a change in output to an exogenous change in the government spending with respect to their respective baselines. Thus, the question is how to calculate fiscal multiplier? There is not the only way to measure fiscal multiplier, and in this paper I will follow the literature to the most widely used way of calculating fiscal multiplier.

[8]A higher unemployment rate z would suggest a higher slackness of the economy.

The i quarter impact fiscal multiplier is:

$$FM(i) = \frac{\Delta Y_i}{\Delta G_i} \tag{3.8}$$

The peak fiscal multiplier is defined as:

$$FMp = \frac{\max_i \Delta Y_i}{\max_i \Delta G_i} \tag{3.9}$$

The cumulative fiscal multiplier, defined as the cumulative change in GDP over the cumulative change in fiscal expenditure at some horizon. The two year cumulative fiscal multiplier is:

$$FM2 = \frac{\sum_{i=1}^{8} \Delta Y_i}{\sum_{i=1}^{8} \Delta G_i} \tag{3.10}$$

The four year cumulative fiscal multiplier is:

$$FM4 = \frac{\sum_{i=1}^{16} \Delta Y_i}{\sum_{i=1}^{16} \Delta G_i} \tag{3.11}$$

3.3 Data Description

We use the data set from Ramey and Zubairy (2014), which starts at 1889 to 2011. The historical series include nominal GDP, government spending, defense news variable, and unemployment rate. All the variables are quarter frequency. This is a new dataset contains more than one hundred years historical observations.

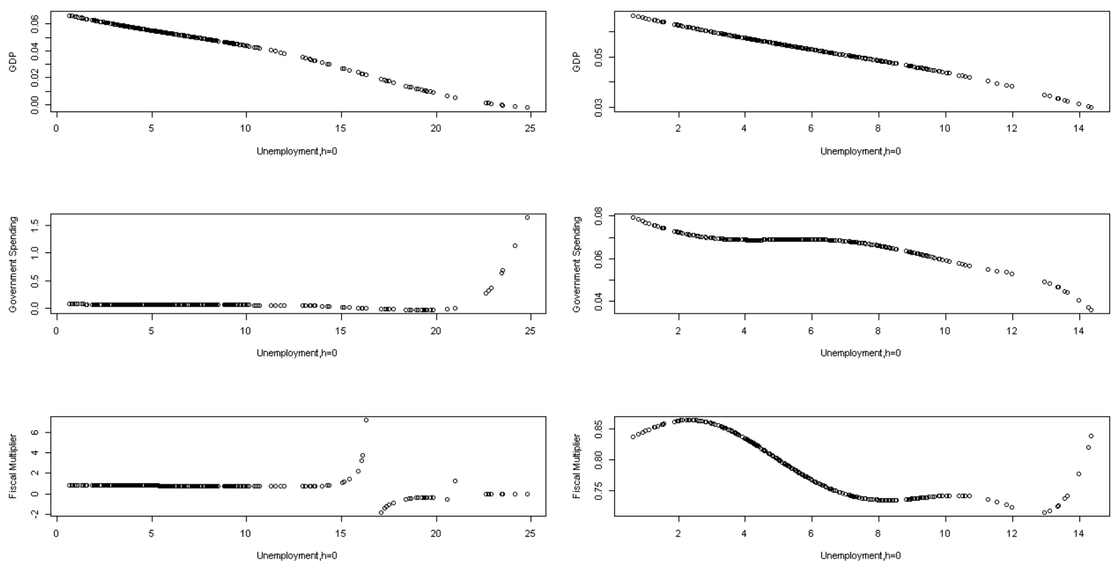
3.4 Empirical Result

In this section, we report the main empirical evidence obtained from aforementioned econometric models.

3.4.1 Generalized State-Dependent Impulse Response Functions

We document the generalized state-dependent impulse response functions and impact fiscal multiplier by seleted periods below. More specifically, Figure 3.1 is the contemporary, Figure 3.2 is 4 quarters, Figure 3.3 is 8 quarters, Figure 3.4 is 12 quarters, Figure 3.5 is 16 quarters, and Figure 3.6 is 20 quarters later effect.

Figure 3.1: Government Spending and GDP Responses to A News Shock, Semiparametric Varying-Coefficient Local Projection, Contemporary.



Compare to the traditional impulse response function calculated by parametric models (either linear or nonlinear), what we report here is different in the sense that the response is not only changing over time, but also a function of some particular state variable (say, unemployment rate in this paper). Moreover, all the following figures show the period by period response from the contemporaneous month to twelve months after, hence the reader may have a clearer picture about how the

Figure 3.2: Government Spending and GDP Responses to A News Shock, Semiparametric Varying-Coefficient Local Projection, 4 Quarters Later.

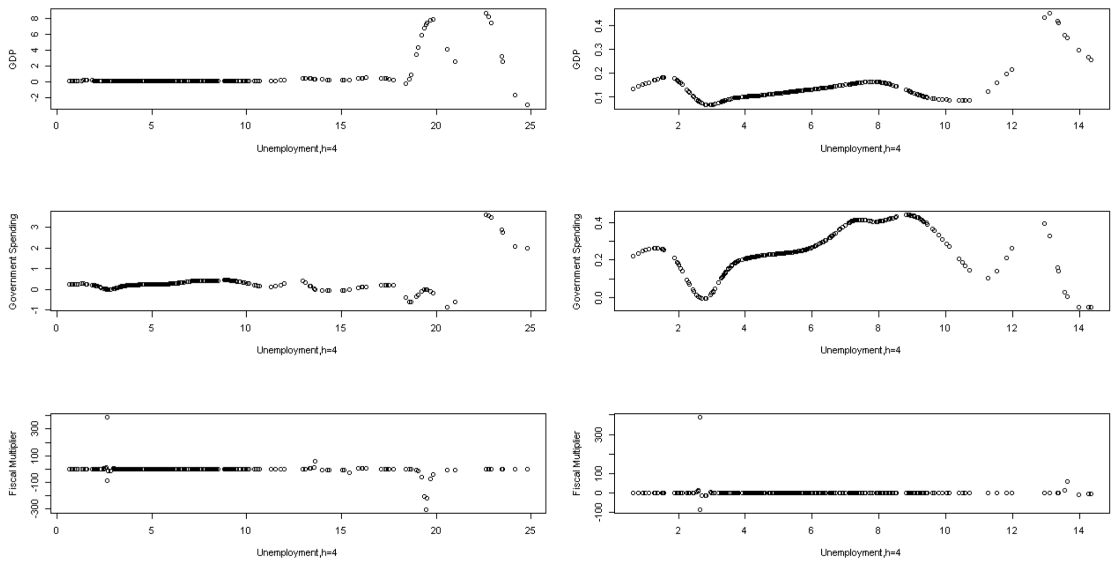


Figure 3.3: Government Spending and GDP Responses to A News Shock, Semiparametric Varying-Coefficient Local Projection, 8 Quarters Later.

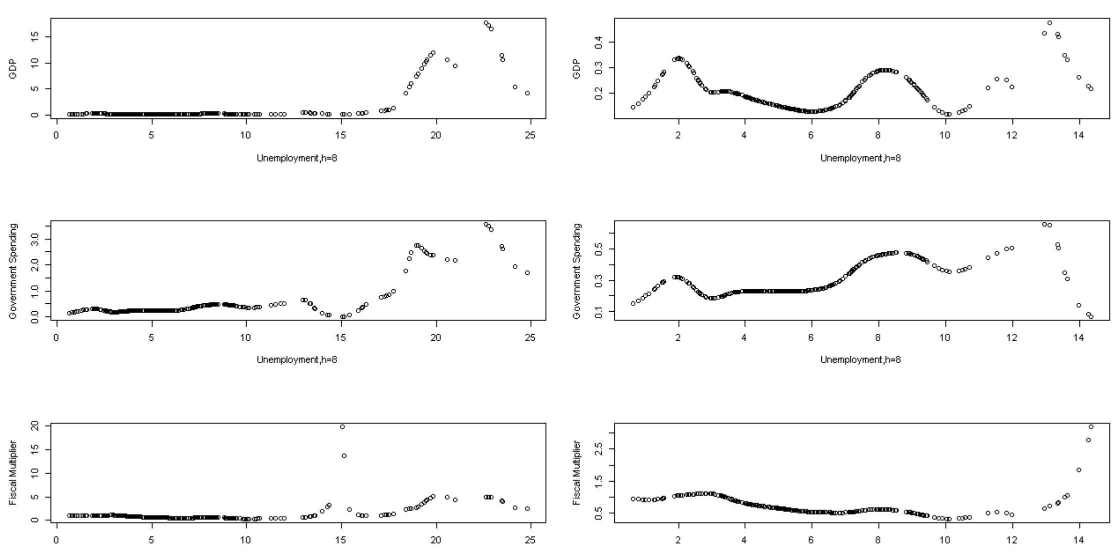


Figure 3.4: Government Spending and GDP Responses to A News Shock, Semiparametric Varying-Coefficient Local Projection, 12 Quarters Later.

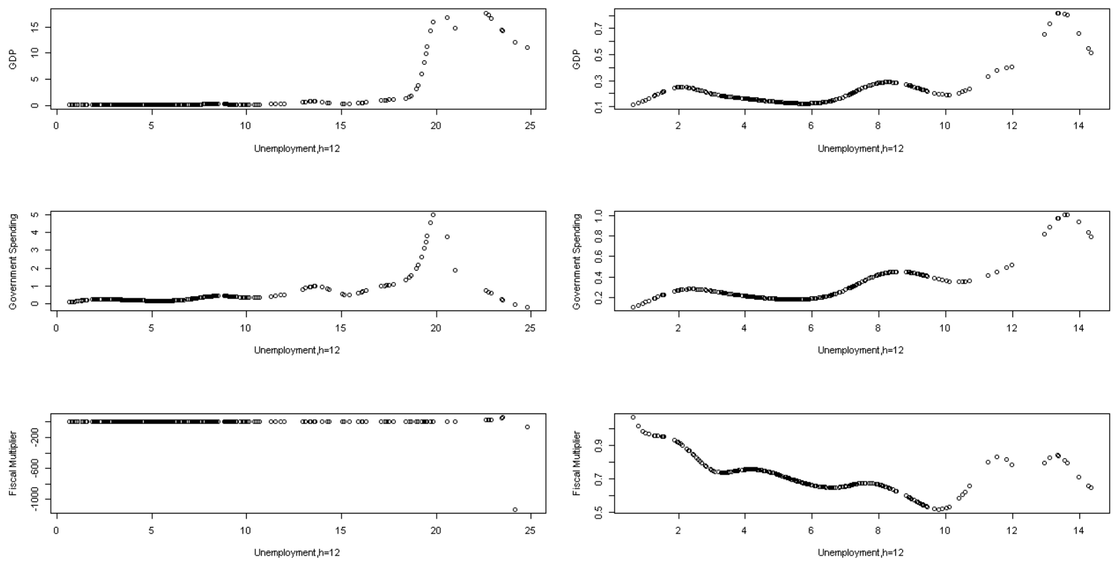


Figure 3.5: Government Spending and GDP Responses to A News Shock, Semiparametric Varying-Coefficient Local Projection, 16 Quarters Later.

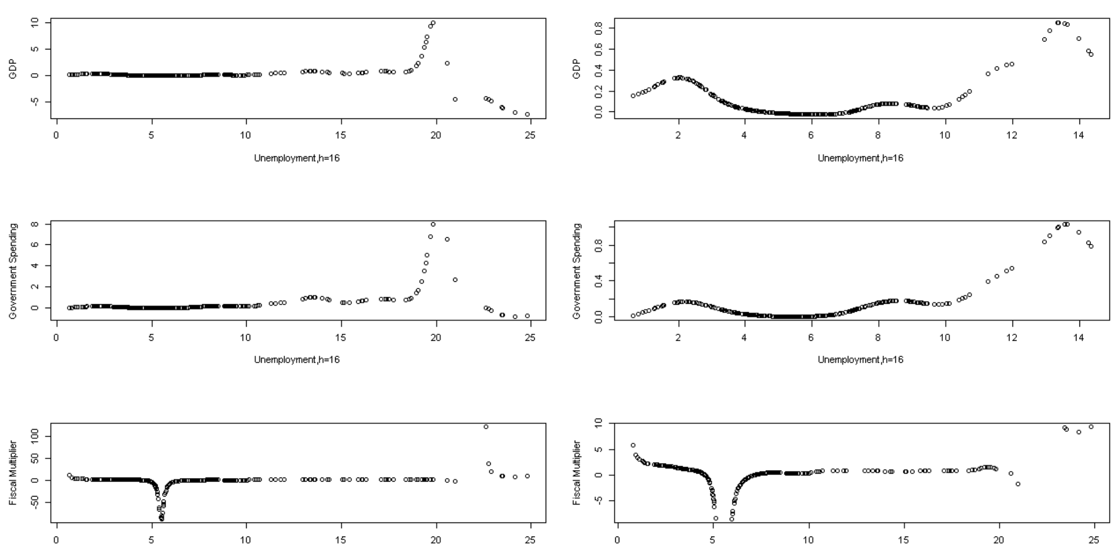
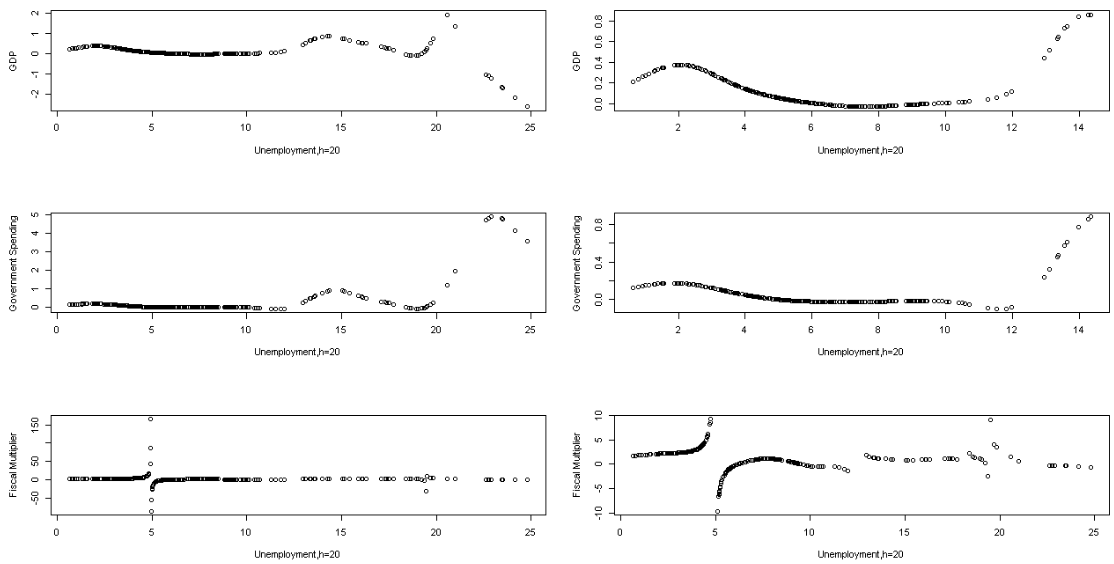


Figure 3.6: Government Spending and GDP Responses to A News Shock, Semiparametric Varying-Coefficient Local Projection, 20 Quarters Later.



effect changes as the unemployment rate varies. Generally speaking, we failed to find any strong systematic pattern of the state dependence effect of fiscal policy on the real economy. Another interesting thing, the shape of the impulse response functions for GDP and government spending is quite similar. Thus, the impact fiscal multiplier is quite flat over the different slack of the economy.

3.4.2 Generalized State-Dependent Fiscal Multiplier

Figure 3.7 is the peak fiscal multiplier, 2 year integral fiscal multiplier, and 4 year fiscal multiplier as a function of the unemployment rate, respectively. All the functions are quite flat, the only exceptions are peak fiscal multiplier and 4 year fiscal multiplier when the unemployment rate is above 18.

One of the most important feature of semiparametric varying-coefficient model in the framework of local projections is data-driven. This gives us lots of freedom

Figure 3.7: Generalized State-Dependent Fiscal Multiplier,
Semiparametric Varying-Coefficient Local Projection.

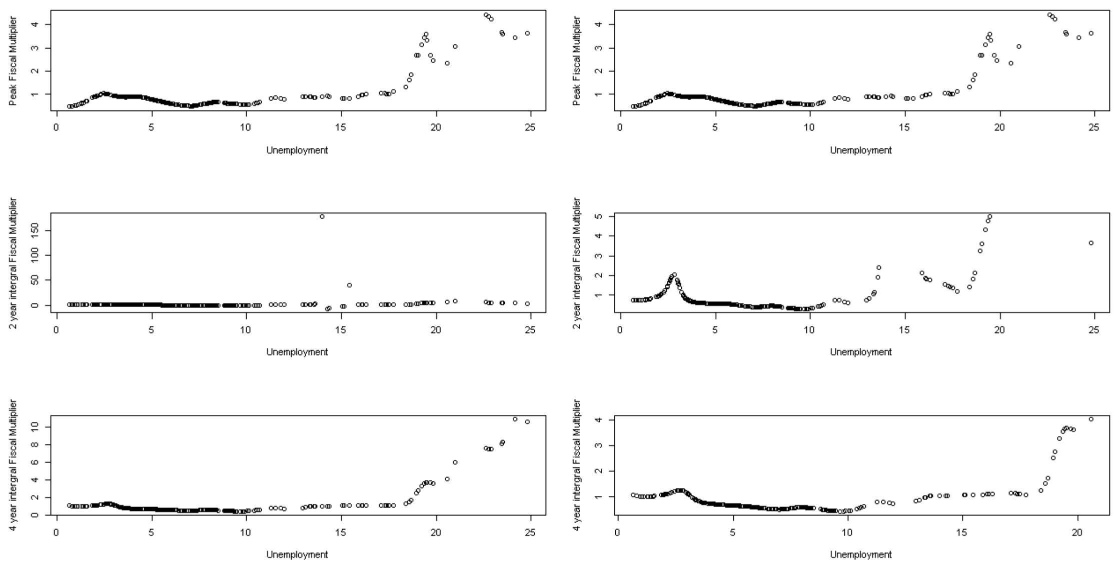


Table 3.1: Estimated Fiscal Multipliers Depending on the Unemployment Rate.

Unemployment Rate	< 6.5	> 6.5	< 5	(5,10)	(10,15)	> 15
Peak Fiscal Multiplier	0.79	0.93	0.88	0.62	0.79	2.29
2 Year Intergral Fiscal Multiplier	0.68	2.28	0.79	0.48	9.31	4.46
4 Year Intergral Fiscal Multiplier	0.68	0.71	0.79	0.48	0.83	2.48

to comprehensively investigate the policy effects. Table 3.1 is a simple illustration to analyze the fiscal multiplier. We document fiscal multipliers for different states of the economy. depending on the range of unemployment rate. For most of cases, fiscal multipliers are slightly below one, and we can not find strong state-dependence effects. The only exception is that when the unemployment rate is above 15. However, since there are few observations belongs to this category, we do not need to worry too much about it.

To sum up, through a flexible methodology, the empirical analysis suggests that fiscal multiplier does not depend on the the unemployment rate. We do not find supporting evidence in favor of old Keynesian channel.

3.5 Conclusion

The Great Recession and subsequent American Recovery and Reinvestment Act fiscal stimulus package have brought fiscal policy to the forefront of policy issues. Is expansionary government spending a good tool to stimulate the economy, or the effect is very small, if any, on consumption and investment? This paper is a comprehensive empirical analysis, aiming to shed new light on the heated debate.

Using the flexible semiparametric varying coefficient method in the framework of local projections, we directly model the fiscal multiplier as a function of various state variables. The paper shows that the U.S. fiscal multiplier is slightly below one and approximately the same, during periods of slack as compared to normal times. Our results suggest that fiscal policy was not necessarily a more powerful tool to stimulate aggregate demand during the "Great Recession".

4. CONCLUSION AND FUTURE RESEARCH

Two empirical works in this dissertation investigate the state-dependent effects of monetary policy and fiscal policy, respectively. We have proposed an empirical framework to study generalized state-dependent impulse response functions, a methodology which should have many applications in other contexts of macroeconomics. This contributes to the literature in the sense that few papers, if any, analyze the continuous state effects of macroeconomic policies.

How financial markets respond to monetary policy has been of interest to both economists and policy makers, more so than ever owing to the Great Recession. In the first work, we find that when the Fed increases the Federal Funds rate, the stock price decreases in normal times, but increases during bubbly episodes. Moreover, the paper captures the effect of an exogenous tightening monetary shock on stock prices as an increasing function of the size of bubbles using a flexible semiparametric varying-coefficient model specification. This paper points out two important transmission channels of monetary policy on asset price: risk premium and asset bubbles, which are often ignored in theoretical models. On the policy side, my empirical analysis suggests that central banks should be cautious about adopting "leaning against bubbles" monetary policies when the bubble size is relatively large.

The conventional interest rate monetary policy analyzed in the first project is ineffective after 2008 due to the zero lower bound constraint. This has brought fiscal policy to the forefront of policy discussions. One key question in the current policy debate is the size of the fiscal multiplier. My second empirical work is directly related to this question. The paper shows that the U.S. fiscal multiplier is slightly below one and approximately the same, during different unemployment rate. The empirical results suggest that fiscal policy is not necessarily a more powerful tool to stimulate aggregate demand during the Great Recession.

The proposed generalized state-dependent method is more informative in measuring macroeconomic policy effects than linear or time-varying methods, and should be appealing to both theoretical macro-economist and policy makers.

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APPENDIX A

LOCAL-LINEAR KERNEL ESTIMATION

We briefly introduce the local-linear kernel method that is used to estimate the semiparametric varying-coefficient model. Following Li and Racine (2007), the local-linear method is based on the following minimization problem:

$$\min \quad \sum_{j=1}^{n}[Y_j - a - (X_j - x)'b]^2 K(\frac{X_j - x}{h}), \tag{A.1}$$

where $K(.)$ is the kernel smooth function, and h is its bandwidth. We can easily use the generalized least square (GLS) to get an estimator, which is proved to be consistent. Moreover, we will use least square cross-validation (LS-CV) to choose an optimal h.

$$\min \quad \frac{1}{n}\sum_{i=1}[Y_i - g_{-i,L}(X_i)]^2 M(X_i) \tag{A.2}$$

where $M(.)$ is weighting function. Compare to the traditional local constant estimator, the local linear method could avert the boundary bias problem.

APPENDIX B

OTHER FIGURES

Figure B.1: S&P 500 Response to Exogenous Monetary Policy (B&C) Shocks, Linear and Two-States Threshold Local Projection

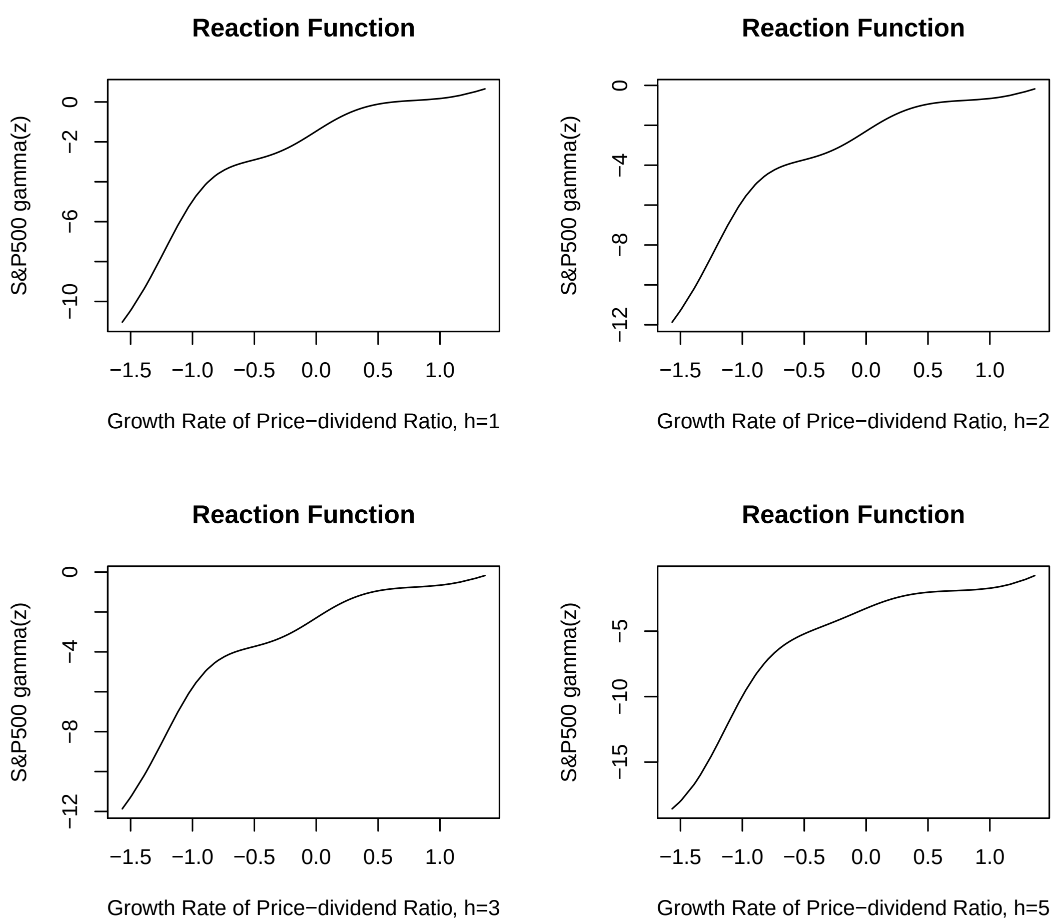

Figure B.2: S&P 500 Response to Exogenous Monetary Policy (B&C) Shocks, Linear and Two-States Threshold Local Projection

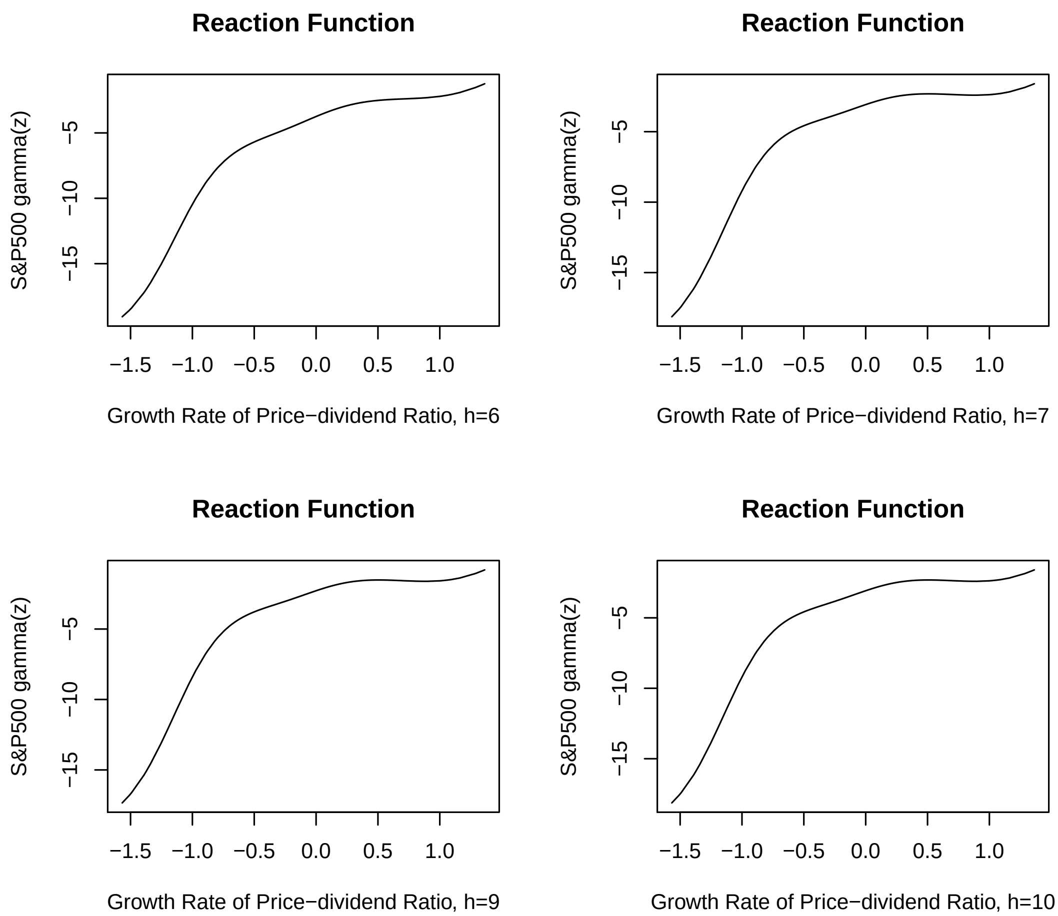

Figure B.3: S&P 500 Response to Exogenous Monetary Policy (B&C) Shocks, Linear and Two-States Threshold Local Projection