ESSAYS ON MONETARY POLICY AND INTERNATIONAL TRADE

A Dissertation

by

HUI-CHU CHIANG

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

DOCTOR OF PHILOSOPHY

May 2008

Major Subject: Economics

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Approved by:

Chair of Committee,	Dennis W. Jansen
Committee Members,	Li Gan
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	David Bessler
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ABSTRACT

Essays on Monetary Policy and International Trade. (May 2008)

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Chair of Advisory Committee: Dr. Dennis W. Jansen

The dissertation consists of three essays. Chapter II examines the asymmetric effects of monetary policy on stock prices by using an unobserved components model with Markov-switching. My results show that monetary policy has negative effects on stock prices, which is consistent with the most recent literature. When the transitory component is in the low volatility state, a contractionary monetary policy significantly reduces stock prices. When the transitory component is in the high volatility state, the negative effect of monetary policy becomes larger, but the difference of the monetary policy effects between two states is not significant. Besides, a contractionary monetary policy will lower the probability of stock prices staying in the low volatility state. Monetary policy also reduces the total volatility of stock prices and the volatility of the transitory component of stock prices.

Chapter III employs the smooth transition autoregressive (STAR) models to investigate the nonlinear effect of monetary policy on stock returns. The change in the Federal funds rate is used as an endogenous measure of monetary policy and the growth rate of industrial production is also considered in the model. My empirical results show

that excess stock returns, the change in the Federal funds rate, and the growth rate of industrial production all can be expressed in the nonlinear STAR models. The estimated coefficients and the impulse response functions show that the effect of monetary policy on excess returns of stock prices is significantly negative and nonlinear. The change in the Federal funds rate has a larger negative effect on excess returns in the extreme low excess returns regime and the effect becomes smaller when the excess returns are greater than the threshold value.

In chapter IV, I use a panel data approach to investigate the impact of exchange rate volatility on bilateral exports of the U.S. to the thirteen major trading partners. I further test the possibility of nonlinear effects of exchange rate volatility on exports by using threshold regression methods for non-dynamic panels with individual-specific fixed effects proposed by Hansen (1999). The results indicate that the effect of exchange rate volatility on bilateral exports is nonlinear. When the relative real GDP per capita of the exporting partner is lower than the threshold value, the response of bilateral U.S. exports to exchange rate volatility is positive. But, exchange rate volatility decreases bilateral exports of the U.S. to the exporting partners when their relative real GDP per capita surpass the threshold value.

DEDICATION

To my parents,

Yin-Fu Chiang and Li-Mei Kuo

for their unconditional love

To my husband, Kuang-Chung

for his eternal encouragement and support,

and to my beloved daughter, Alice

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I would like to express my appreciation to all those who have provided assistance and encouragement to this research. Special thanks to my committee chair, Dr. Dennis Jansen, for his dedicated attention, guidance, and encouragement to this research. I also would like to thank my other committee members, Dr. Li Gan, Dr. Paula Hernandez-Verme, and Dr. David Bessler, for their support and comments that were very helpful during the final stage of my research. I am also thankful to the Department of Economics at Texas A&M University for providing the financial support for my studies.

I am also grateful to my family and friends for all of their unconditional support and love. I am especially grateful to my husband, Kuang-Chung, for giving me everything that I need. I can study abroad and have finished my Ph.D. program all because of you.

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CHAPTER I

INTRODUCTION

In this dissertation, I investigate the nonlinear relationship between economics variables in the field of monetary economics and international trade. Chapter II and Chapter III examine the nonlinear effect of monetary policy on stock return and stock prices by using two different kinds of nonlinear models. Chapter IV discusses the threshold effect of real exchange rate volatility on bilateral exports.

In the recent years, the global economy has experienced many times of financial crashes and booms. Economists are paying more attention on the relationship between monetary policy and financial market and trying to find out if monetary policy can affect the financial market. A large number of studies have tried to investigate the effects of monetary policy on stock returns from every perspective. They use different kind of monetary policy variable, for example, money aggregate data (Pesando(1974), Rogalski and Vinso(1977)), the changes in market interest rate or official rate (Patelis (1997), Perez-Quiros and Timmermann (2000)), or extracting the unexpected monetary policy shocks, such us the orthogonalized innovations from a vector autoregressive model (Thorbecke (1997), Chen(2005)), etc. Some literature use a variety of empirical techniques, for instance, the vector autoregression estimation, generalized method of moments estimation, or an event study methodology. Most of papers discuss the linear response of stock returns to monetary policy and find a negative effect of monetary

This thesis follows the style of the Journal of Monetary Economics.

policy on stock returns, but not many researchers focus on the possible nonlinear relationship between monetary policy and stock returns, even though it might happen in the theoretical point of view.

According to the theoretical model, information asymmetry might exist in financial markets, and then agents may behave as if they were constrained. The financial constraint problem could be more serious in the bad economic environment. This implies that the monetary policy might have asymmetric effects on financial market and the asymmetry effects might be determined by the situation of stock market, the state of economy, or monetary policy itself. However, only a few studies examine the asymmetric effects of monetary policy on stock market and they just use the simple dummy variables in their equations, except Chen (2007). Chen investigates the asymmetric monetary policy effects on stock returns by using Markov-switching models. He finds that monetary policy has larger effects on stock returns in bear market and a contractionary monetary policy leads to a higher probability of switching to the bear market regime.

The motivation of chapter II and chapter III is to discuss the effect of monetary policy on stock market by using nonlinear models. According to the results of Summers (1986), Fama and French (1988), Kim and Kim (1996), the unobserved-components model with Markov-switching is a good model to illustrate stock prices. So, in chapter II, I attempt to investigate the asymmetric effects of monetary policy on stock market by using an unobserved-components model with Markov-switching (UC-MS model). I augment UC-MS model with a monetary policy variable and assume that monetary policy only influences the transitory component of stock prices. First, I estimate UC-MS model with no monetary policy as a benchmark model. Then, the model is augmented with monetary policy variable for the purpose of investigating the effects of monetary policy on stock prices. I also estimate a third model which allows the transition probability to be time-varying, which depends on monetary policy shocks.

Chapter III uses smooth transition autoregressive (STAR) models to investigate the nonlinear effect of monetary policy on stock returns. I consider three important variables, stock returns, the change in the Federal funds rate, and the growth rate of output. Since three variables are all endogenously determined, the STAR models are constructed for each variable. Besides, stock returns, the change in the Federal funds rate, and the growth rate of output are all allowed to be the possible threshold variable which controls for the nonlinear dynamics of models. By appropriately choosing the best threshold variable for the model of each variable and estimating the nonlinear models for them, the nonlinear relationship among excess stock returns, monetary policy, and output growth can be investigated. Finally, the nonlinear impulse response functions are calculated in order to understand how they affect to each other.

Chapter IV investigates nonlinear effects of exchange rate volatility on exports. In the previous literature, the effect of real exchange rate volatility on exports has been fully discussed by using time series data, but the conclusion is still mixed, especially using the bilateral exports data. From the theoretical point of view, De Grauwe (1988) argues that the impact of exchange rate volatility on exports depends on the degree of risk aversion. If the trader is less risk averse, the income effect might be greater than the substitution effect when exchange rate volatility increases, and will increases exports.

I attempt to reexamine the effects of real exchange rate volatility on bilateral exports by using panel data approach. The data of real bilateral export volume from U.S. to thirteen major trading partners are used. I further test the possibility of nonlinear effects of exchange rate volatility on exports by using threshold regression methods for non-dynamic panels with individual-specific fixed effects proposed by Hansen (1999). Referring to the most empirical papers, the bilateral exchange rate volatility is measured by using moving sample standard deviation method and the conditional standard deviation from a GARCH (1,1) model. In order to check the robustness of conclusion, the model is estimated again for top 30 major exporting partners of the United States.

CHAPTER II

THE EFFECT OF MONETARY POLICY ON STOCK PRICES: AN UNOBSERVED-COMPONENTS MODEL WITH MARKOV-SWITCHING

2.1 Introduction

After the collapse of the Japanese and U.S. asset price bubbles, the relationship between monetary policy and asset prices has brought people's new attention. One of the important issues is the role of asset prices in the monetary transmission mechanism. The transmission mechanism of monetary policy usually comes through the stock market by changing the values of private portfolios (the wealth effect) and the cost of capital, thus, in turn affects the real economy. So, the purpose of this paper is trying to understand the role of asset prices in the monetary transmission by estimating the effects of monetary policy on stock prices.

In the early empirical studies, they usually estimate the effects of monetary policy on asset prices by using money aggregate data as the monetary policy variable. However, the results are not consistent among all the research. For example, Pesando (1974) uses linear regressions and finds that no impacts of changes in the money supply on stock prices. Rogalski and Vinso (1977) estimate cross correlations between money supply and stock prices and conclude that there is no significant forecasting power of changes in money on stock prices. But, Homa and Jaffee (1971) find that expansionary policy increases stock prices by using linear regressions. Recently, after Bernanke and

Blinder (1992) show that Federal funds rate is a good indicator of monetary policy actions, economists re-estimate the link between monetary policy and stock market. Most results agree that monetary policy helps to explain the stock prices or returns and the effect of monetary policy on stock market is negative. Thorbecke (1997) uses Vector Autoregressive model and concludes that a contractionary monetary policy decreases stock returns. By using event-study approach, Rigobon and Sack (2004) find an increase in the short-term interest rate has a negative impact on stock prices; Bernanke and Kuttner (2005) conclude that unexpected cut in the Federal funds rate would lead an increase in stock prices.

According to the theoretical model, information asymmetry might exist in financial markets, and then agents may behave as if they were constrained. The financial constraint problem could be more serious in the bear markets. This implies that monetary policy might have asymmetric effects on financial market between different financial states. However, only some of the previous studies examine the asymmetric effects of monetary policy on stock market and they just use the simple dummy variables in their equations, except Chen (2007). Chen investigates the asymmetric monetary policy effects on stock returns by using Markov-switching models. He finds that monetary policy has larger effects on stock returns in bear market and a contractionary monetary policy leads to a higher probability of switching to the bear-market regime.

For modeling stock prices, Summers (1986) proposed an unobserved-components model (UC model). Summers decomposes stock prices into a stochastic trend component and a stationary transitory component and finds that the stationary transitory component of prices accounts for a substantial fraction of the variation of returns. Fama and French (1988) conclude that the existence of stationary transitory components of stock prices is more significant for the portfolio of small firms than for the portfolio of large firms. Kim and Kim (1996) then add the Markov-switching (MS) method into an UC model and use data from 1951:1 to 1992:12. They find that the UC-MS model describes the pattern of stock prices well and can capture the quick volatility reverting of stock returns to its normal level after the crash.

This paper attempts to investigate the asymmetric effects of monetary policy on stock market by using an unobserved-components model with Markov-switching (UC-MS model) from Kim and Kim (1996) instead of the usual linear model in most previous papers. I augment UC-MS model with a monetary policy variable to discuss the possibility of nonlinear effects of monetary policy on stock prices and assume that monetary policy only influences the transitory component of stock prices. First, I estimate the UC-MS model with no monetary policy as a benchmark model. Then, the UC-MS model with monetary policy variable is estimated to understand the asymmetric effects of monetary policy on stock prices. In addition, I also estimate a third model which allows the transition probability to be time-varying, where time variation depends on monetary policy shocks.

My results show that monetary policy has negative effects on stock prices, which is consistent with the most recent literature. When the transitory component is in the low volatility state, a contractionary monetary policy reduces stock prices and the effect is significant. When the transitory component is in the high volatility state, the negative effect of monetary policy becomes larger, but the difference of the monetary policy effects between two states is not significant. Besides, monetary policy can affect the dynamics of switching between low volatility and high volatility state. A contractionary monetary policy will lower the probability of stock prices staying in the low volatility state. Monetary policy also reduces the total volatility of stock prices and the volatility of the transitory component.

The remainder of the chapter proceeds as follows. Section 2 describes the empirical model to be estimated. Section 3 contains the data information and how to measure the monetary policy variable. Section 4 presents the empirical results. Section 5 is the conclusion.

2.2 Empirical Model

Consider an unobserved components model with Markov-switching heteroscedasticity (UC-MS model) from Kim and Kim (1996):

$$p_t = p_t^P + p_t^T , \tag{2.1}$$

$$p_t^P = \mu + p_{t-1}^P + v_t , \tag{2.2}$$

$$p_t^T = \phi_1 p_{t-1}^T + \phi_2 p_{t-2}^T + e_t , \tag{2.3}$$

$$v_t \sim N(0, \sigma^2_{v,S_t}) , \quad \sigma^2_{v,S_t} = (1 - S_{vt})\sigma^2_{v0} + S_{vt}\sigma^2_{v1} , \quad \sigma^2_{v1} > \sigma^2_{v0} , \tag{2.4}$$

$$e_t \sim N(0, \sigma^2_{e,S_t}) , \quad \sigma^2_{e,S_t} = (1 - S_{et})\sigma^2_{e0} + S_{et}\sigma^2_{e1} , \quad \sigma^2_{e1} > \sigma^2_{e0} , \tag{2.5}$$

where p_t is the log of real stock prices, which is decomposed into a stochastic trend component p_t^P and a stationary transitory component p_t^T. Equation (2.2) means that the stochastic trend component p_t^P is specified as a random walk with a drift term μ. In equation (2.3), the transitory component is assumed to be a stationary autoregressive process. S_{vt} and S_{et} are discrete-valued, unobserved first-order Markov-switching variables which equal either 0 or 1. v_t, e_t are the innovations to p_t^P and p_t^T, which are assumed to have Markov-switching variances in the form of equation (2.4) and (2.5). We assume that the variances of v_t and e_t are larger in state 1 than in state 0.

In order to investigate the asymmetric effects of monetary policy on stock prices, I modify the model of Kim and Kim (1996). Equation (2.3) is augmented as follows:

$$p_t^T = \phi_1 p_{t-1}^T + \phi_2 p_{t-2}^T + \beta_0 x_{t-1} + \beta_1 x_{t-1} S_{et} + e_t \,, \qquad (2.3')$$

where x_{t-1} is the monetary policy variable at time t-1. I assume that monetary policy has no effect on permanent stock prices. It only changes the transitory component of stock prices. The effect of monetary policy might be asymmetric, which depends on the state of innovation e_t of the transitory component. For example, when S_{et} equals to 0, the effects of x_{t-1} on p_t^T is β_0; when S_{et} equals to 1, the effects of x_{t-1} on p_t^T is $\beta_0 + \beta_1$.

For the unobserved state variable S_{vt} and S_{et}, I first assume that they evolve independently of each other according to the following transition probabilities:

$$P(S_{v,t} = 0 \mid S_{v,t-1} = 0) = p_{00}^v \,,\; P(S_{v,t} = 1 \mid S_{v,t-1} = 1) = p_{11}^v \qquad (2.6)$$

$$P(S_{e,t} = 0 \mid S_{e,t-1} = 0) = p_{00}^{e}, \; P(S_{e,t} = 1 \mid S_{e,t-1} = 1) = p_{11}^{e} \tag{2.7}$$

p_{00}^{v} is the probability of the trend component of stock prices moving from state 0 at time t-1 to state 0 at time t. The stochastic processes of S_{vt} and S_{et} are assumed to be fixed, not determined by any other exogenous or predetermined variables.

Then, I modify equation (2.7) to relax the assumption of fixed transition probability and allow the transition probabilities of the regime-switching process of S_{et} to be time-varying, where time variation depends on monetary policy shocks. The functions of the transition probabilities are then specified as follows:

$$P(S_{e,t} = 0 \mid S_{e,t-1} = 0) = p_{00}^{e}(x_{t-1}) = \frac{\exp(c_0 + a_0 * x_{t-1})}{1 + \exp(c_0 + a_0 * x_{t-1})} \tag{2.8}$$

$$P(S_{e,t} = 1 \mid S_{e,t-1} = 1) = p_{11}^{e}(x_{t-1}) = \frac{\exp(c_1 + a_1 * x_{t-1})}{1 + \exp(c_1 + a_1 * x_{t-1})} \tag{2.9}$$

The estimates of a_0 and a_1 indicate how monetary policy affects the shifts between high volatility and low volatility state of the transitory component. For example, $a_0 < 0$ implies that a contractionary monetary policy $x_{t-1} < 0$ makes the low volatility state more possible to turn into the high volatility state. In contrast, $a_0 > 0$ indicates that a contractionary monetary policy makes the transitory component of stock prices more likely to stay in the low volatility state.

Before estimating the model, I need to rewrite it as a state-space with Markov switching representation which consists of a measurement equation and a transition equation. The measurement equation is an equation that describes the relation between observed variables (data) and unobserved state variables. The transition equation is an equation that describes the dynamics of the state variables. I take first difference of equation (2.1) and substitute equation (2.2) into it to get the measurement equation

$$r_t = \mu + \begin{pmatrix} 1 & -1 \end{pmatrix} \begin{pmatrix} p_t^T \\ p_{t-1}^T \end{pmatrix} + v_t \tag{2.10}$$

where $r_t = p_t - p_{t-1}$. The transition equation (11) is obtained from equation (3).

$$\begin{pmatrix} p_t^T \\ p_{t-1}^T \end{pmatrix} = \begin{pmatrix} \phi_1 & \phi_2 \\ 1 & 0 \end{pmatrix} \begin{pmatrix} p_{t-1}^T \\ p_{t-2}^T \end{pmatrix} + \begin{pmatrix} \beta_0 x_{t-1} + \beta_1 x_{t-1} S_{et} \\ 0 \end{pmatrix} + \begin{pmatrix} e_t \\ 0 \end{pmatrix} \tag{2.11}$$

Next, these two equations are estimated by using Kim's (1994) basic filter which is a combination of the Kalman filter and Hamilton filter, along with appropriate approximations to get the maximum likelihood estimates { p_{00}^v, p_{11}^v, p_{00}^e, p_{11}^e, σ_{v0}, σ_{v1}, σ_{e0}, σ_{e1}, ϕ_1, ϕ_2, μ, β_0, β_1 }.[1]

2.3 Data

The data frequency is monthly. The stock prices p_t are measured by the log of real stock prices index of New York Stock Exchange (NYSE) which is deflated by the CPI which is deflated by the CPI. For the monetary policy variable x_t, I use monetary policy shocks, the orthogonalized innovations from the standard Vector Autoregression

[1] See Kim and Nelson (1999) for more detail of estimation and applications of state-space with Markov switching models.

(VAR) Model proposed by Christiano et al. (1999), to avoid the endogeneity problem. Federal funds rates are used to be the monetary policy instrument.

The VAR model for extracting monetary policy shocks is

$$Z_t = A_0 + A_1 Z_{t-1} + A_2 Z_{t-2} + \ldots + A_q Z_{t-q} + u_t \tag{2.12}$$

where Z_t is { Y_t , CPI_t , $PCOM_t$, FF_t , NBR_t , TR_t }. Y_t is the log of industrial production, CPI_t is the log of consumer price index, $PCOM_t$ is the log of commodity prices, FF_t is the Federal funds rate, NBR_t is the log of non-borrowed reserves, and TR_t is the log of total reserves. Those variables are suggested by Christiano et al. (1999) and the order of the variables in the vector Z_t is the same as the order in which they are listed above. u_t is serially uncorrelated and has variance-covariance matrix V . The VAR disturbances are assumed to be related to the underlying economic shocks, ε_t , by

$$u_t = C\varepsilon_t \tag{2.13}$$

where C is lower triangular and ε_t has covariance matrix equal to the identity matrix. The VAR is estimated over the period from 1959:01 to 2005:12. After checking the SIC, two lags of each variable are used in the VAR model and the residuals u_t and the variance-covariance matrix V can be obtained after estimation. Then, I can calculate C by the relationship $V = CC'$ and have the underlying economic shocks ε_t by the equation $\varepsilon_t = C^{-1}u_t$. The orthogonalized residuals of Federal funds rate, the fourth element ε_t , is used as the monetary policy variable x_t

All data are collected from Federal Reserve Economic Data and the Center for Research in Security Prices (CRSP) database. The original data are from 1959:01 to 2005:12. Due to the lost of two observations for extracting Federal funds rate shocks, the sample period for the model is from1959:03 to 2005:12.

2.4 Estimation Results

2.4.1 The UC-MS Model without Monetary Policy

Before estimating the monetary policy effects on stock prices, I estimate the model without monetary policy in equation (2.1)-(2.5) as a benchmark model, i.e. I estimate Kim and Kim (1996)'s model by using real stock prices index of NYSE for the period 1959:3 to 2005:12. In the estimating process, σ_{e0} falls on the boundary value zero and makes difficulties in inverting the information matrix to get the standard errors for other parameters. Thus, I impose σ_{e0} to be zero and continue the optimization with respect to other parameters. To make sure that $\sigma_{e0} = 0$ has the maximum log likelihood value, I restrict σ_{e0} with different values and re-estimate the model to check whether $\sigma_{e0} = 0$ is the best estimate. Figure 2.1 shows the graph log likelihood value with σ_{e0}. The result shows that the log likelihood value decreases when σ_{e0} increases. Therefore, $\sigma_{e0} = 0$ is the maximum likelihood estimates. Kim and Kim (1996) have the same result $\sigma_{e0} = 0$ for S&P 500 stock price index from 1952:1 to 1992.12.

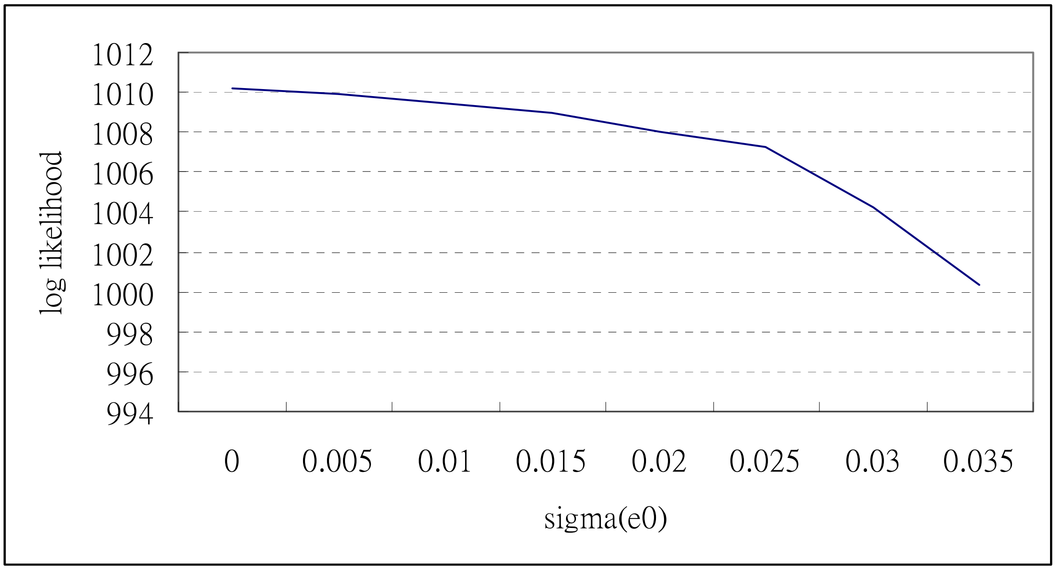

Figure 2.1. Log likelihood value with different σ_{e0}

Estimates of the model with no monetary policy, named Model 1, are reported in the second column in Table 2.1. State 0 represents the low-volatility state and state 1 represents the high-volatility state for both the trend component and the transitory component. The standard error of the trend component shocks is 0.022 in the low-volatility state and is 0.041 in the high-volatility state. The estimates of transition probability of the trend component (p_{00}^{v} and p_{11}^{v}) indicate that the probability of the trend component of stock prices stay in state 0 from time t-1 to time t is 0.990 and stay in state 1 from time t-1 to time t is 0.969. The trend component has high probability to stay in the same state. For the estimates of the parameters associated with the transitory component, the standard error of shocks σ_e is zero in the low-volatility state, while it is large and significant for the high-volatility state. The result is similar to the estimates of Kim and Kim (1996) and they explain that the transitory component is either on or off over the time period. The estimates of transition probability of the transitory component (p_{00}^{e} and p_{11}^{e}) are 0.984 and 0.590, respectively. The expected durations of the low-volatility state is 1/(1-0.984) = 62.5 months, but those of the high-volatility state is only 1/(1-0.590) = 2.44 months. It indicates that the low-volatility state dominates the high-volatility state.

Table 2.1 Maximum likelihood estimates of the unobserved-components model with Markov-switching heteroscedasticity: 1959:01 ~ 2005:12

	Model 1	Model 2	Model 3	
	No monetary policy	Has monetary policy	Time-varying transitory probability	
p_{00}^{v}	0.969 (0.021)	0.968 (0.020)	0.968 (0.021)	
p_{11}^{v}	0.990 (0.007)	0.990 (0.007)	0.990 (0.007)	
σ_{v0}	0.022 (0.002)	0.022 (0.002)	0.022 (0.002)	
σ_{v1}	0.041 (0.002)	0.041 (0.002)	0.040 (0.022)	
p_{00}^{e}	0.984 (0.012)	0.986 (0.009)	c_0 = 4.539 (0.832)	a_0 = -3.283 (1.309)
p_{11}^{e}	0.590 (0.171)	0.652 (0.172)	0.523 (0.358)	
σ_{e0}	0 (-)	0 (-)	0 (-)	
σ_{e1}	0.0090 (0.026)	0.086 (0.025)	0.079 (0.018)	
ϕ_1	1.058 (0.205)	1.292 (0.154)	1.199 (0.156)	
ϕ_2	-0.250 (0.175)	-0.367 (0.139)	-0.313 (0.141)	
μ	0.006 (0.001)	0.006 (0.001)	0.006 (0.001)	
β_0		-0.012 (0.004)	-0.012 (0.005)	
β_1		-0.017 (0.040)	-0.002 (0.017)	
Log likelihood value	-1010.187	1015.688	1019.6202	

Note: Figures in parentheses are approximate standard errors.

Figure 2.2a shows the stock prices and the estimated trend component of stock prices of Model 1. They match well except some periods of time. Figure 2.2b shows the estimated transitory component and the filtered probabilities of the trend component and the transitory component, which are the estimated probabilities that $S_{vt} = 1$ at time t and $S_{et} = 1$ at time t, respectively. S_{et} is generally equal to zero, but it switches to one occasionally. The transitory component usually fluctuates around zero, but it reduces a lot in some periods and the timing is almost the same as $P(S_{et} = 1 | t)$ jumps up. It indicates that when the transitory component of stock price decreases more, it is more likely to be in the high volatility state. I find that every period that the transitory component drops a lot and $P(S_{et} = 1 | t)$ is high is close to one of the crashes in the stock market, such as oil crisis in 1973, black Monday in 1987, Asian financial crisis in 1997, and Russian financial crisis in 1998 etc.

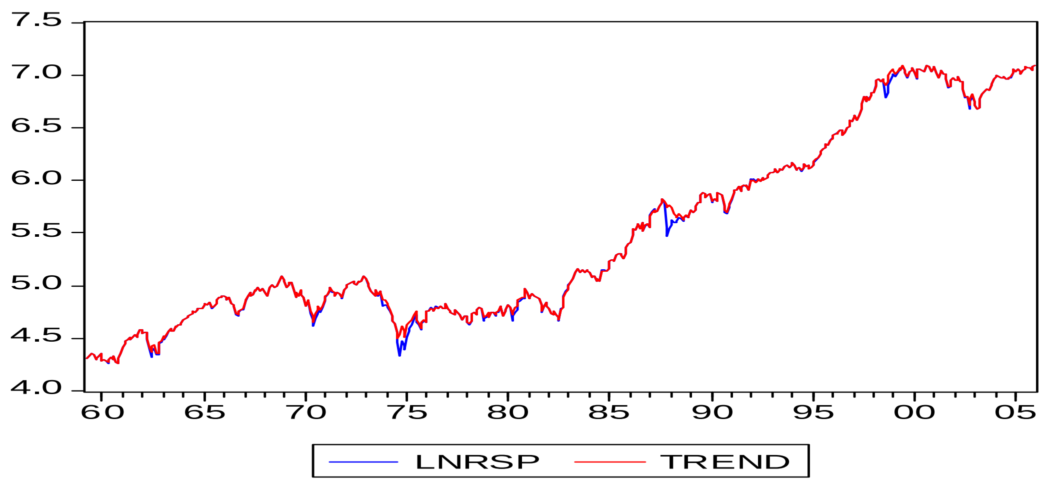

Figure 2.2a: Stock prices and the estimated trend component for Model 1

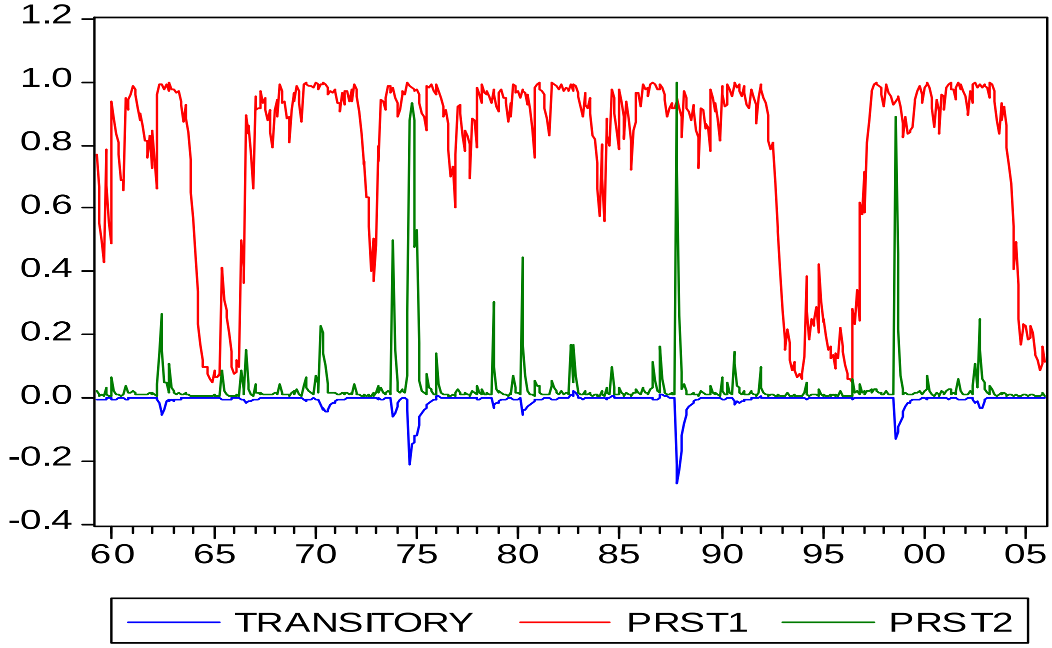

Figure 2.2b. The estimated transitory component and the filtered probabilities of the trend component and the transitory component for Model 1

2.4.2 The UC-MS Model with Monetary Policy

Next, I consider the UC-MS model with monetary policy, named Model 2, to investigate the effects of monetary policy on the transitory component of stock prices and compare the results with the benchmark model. The estimation results of Model 2 are reported in the third column in Table 2.1. The standard errors and transition probabilities of the trend component for both volatility states are almost the same as in Model 1. For the transitory component, the standard error of shocks σ_{e0} and transition probabilities p_{00}^{e} in the low-volatility state are similar to Model 1, but σ_{e1} and p_{11}^{e} in the high-volatility state are a little different. Compare to the results of Model 1, after including monetary policy, the standard error of shocks σ_{e1} decreases a little bit and the probability of staying in the high volatility state in the next period p_{11}^{e} increases. It indicates that monetary policy helps to lower the volatility when stock market is in the high volatility state, but extends the duration of staying in the high volatility state.

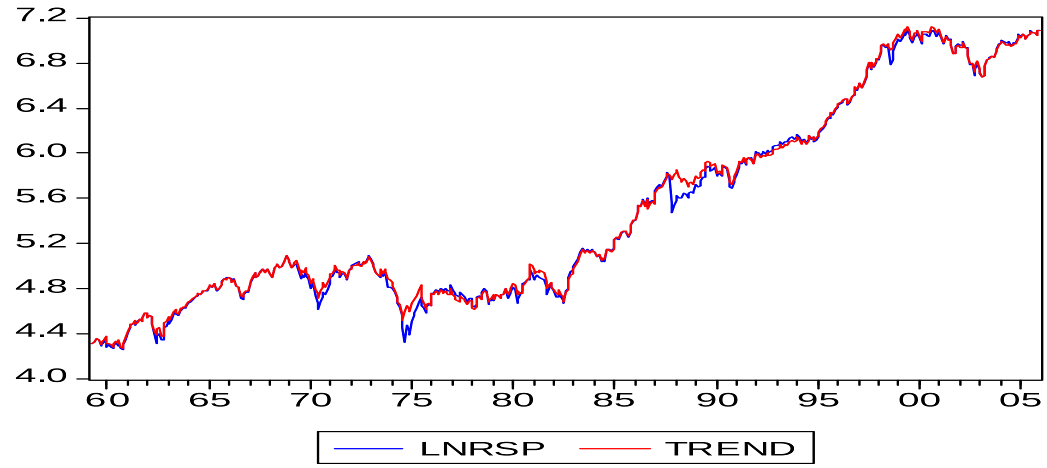

Figure 2.3a. Stock prices and the estimated trend component for Model 2

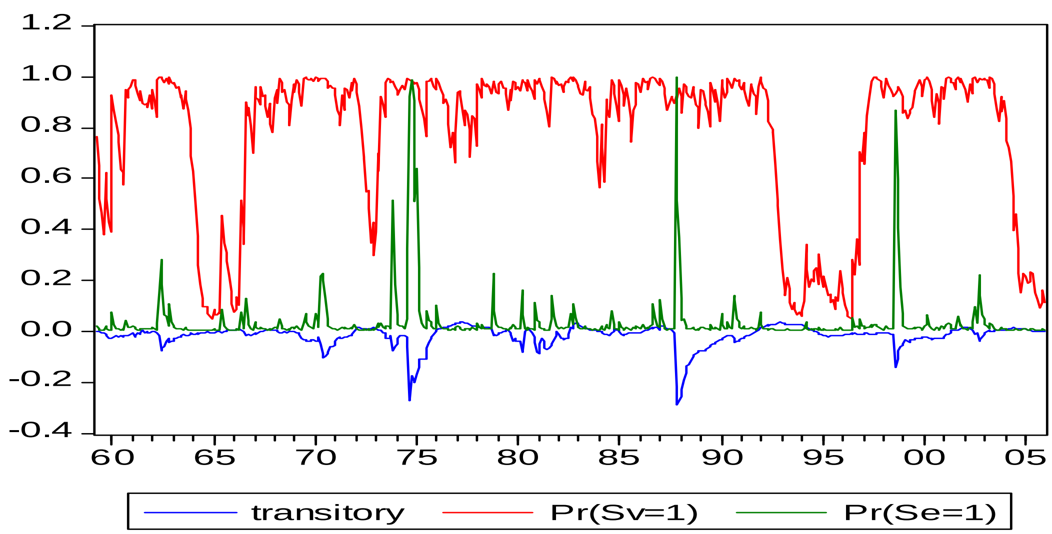

Figure 2.3b. The estimated transitory component and the filtered probabilities of the trend component and the transitory component for Model 2

Figure 2.3a shows the stock prices and the estimated trend component of stock prices of Model 2. The trend component of stock prices is still close to the stock prices, but not that match as in the Model 1. Figure 2.3b shows the estimated transitory component and the filtered probabilities of the trend component and the transitory component of Model 2. Monetary policy increases the fluctuation of the transitory component. The change can be clearly seen when I graph the transitory components of Model 1 and Model 2 together in Figure 2.4.

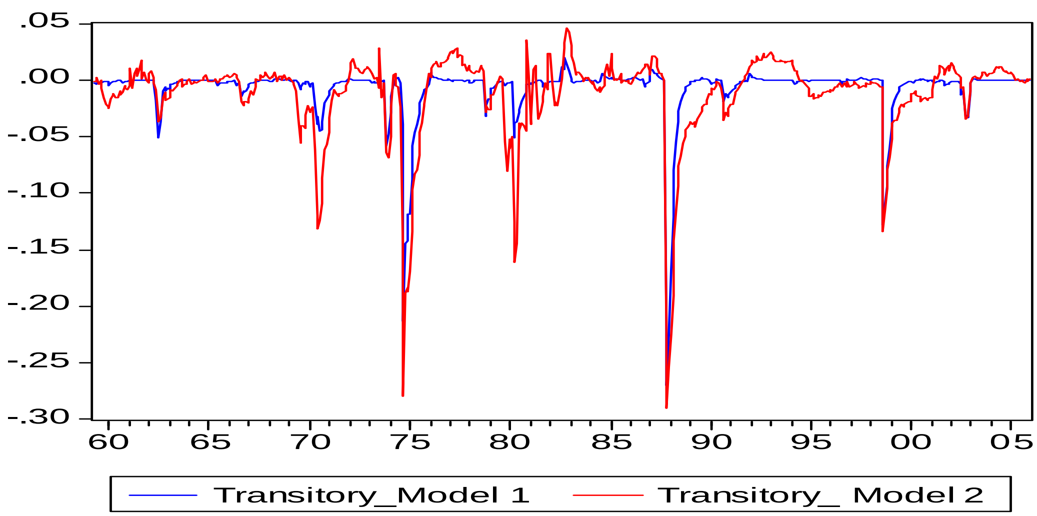

Figure 2.4. The transitory components of Model 1 and Model 2

Moreover, the monetary policy changes the volatility of stock prices, too. Figure 2.5a shows the total volatility of Model 1 and Model 2. Figure 2.5b shows the difference of total volatility between Model 1 and Model 2, which is calculated as total volatility of Model 1 minus the total volatility of Model 2. From Figure 2.5b, the difference of total volatility is usually above zero and has larger positive values on April

1980 and December 1987, even though it is negative for some data points. Figure 2.6a shows the volatility of the transitory component of Model 1 and Model 2. Figure 2.6b shows the difference of volatility of the transitory component between Model 1 and Model 2, which is calculated as volatility of the transitory component of Model 1 minus that of Model 2. The difference is also positive most of the time. Therefore, my results indicate that monetary policy can reduce the volatility of stock market.

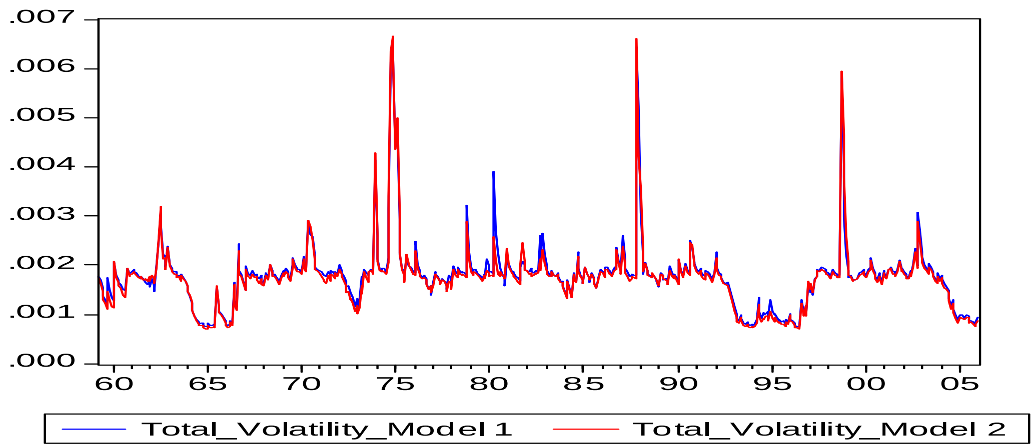

Figure 2.5a. Total volatility of Model 1 and Model 2

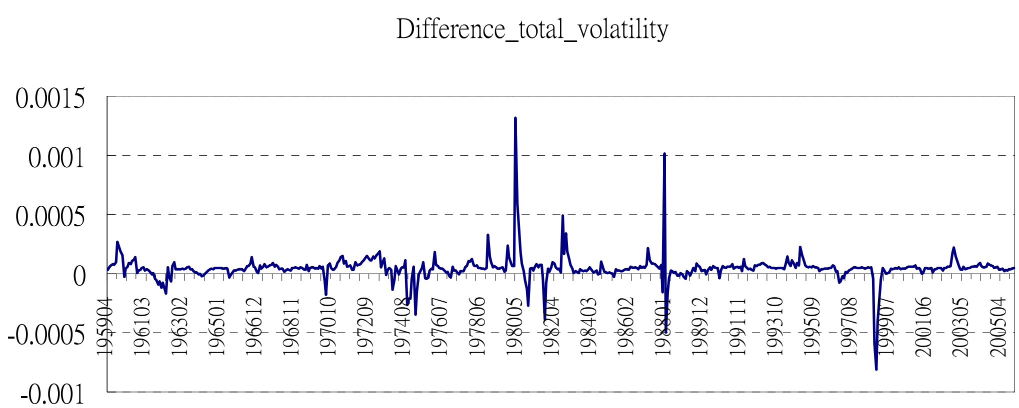

Figure 2.5b. Difference of total volatility between Model 1 and Model 2

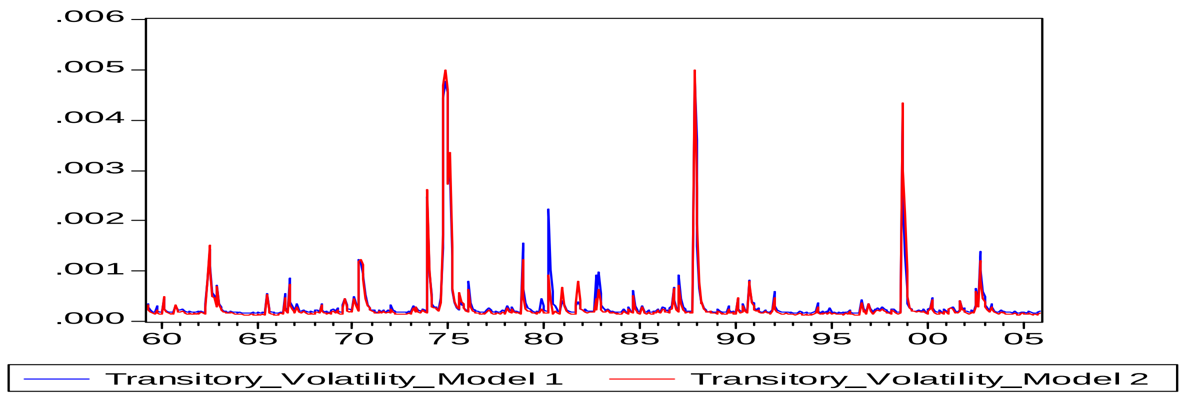

Figure 2.6a. Volatilities of the transitory component of Model 1 and Model 2

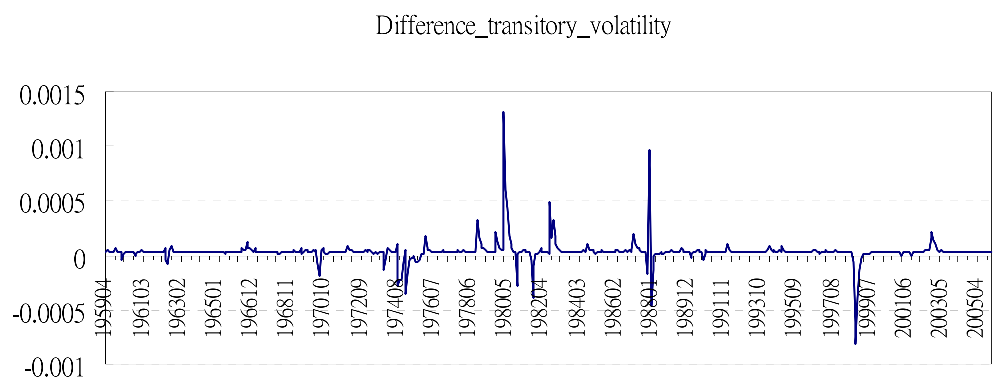

Figure 2.6b. Difference of volatility of the transitory component between Model 1 and Model 2

From the coefficients of Federal funds rate shocks in the third column in Table 2.1, β_0 shows that monetary policy shocks have significant negative effects on the transitory component of stock prices in the low-volatility regime, 1% increase in the Federal funds rate shocks decreases stock prices by 0.012%. The estimate of β_1 shows that monetary policy shocks have larger negative effects on stock prices (β_0 + β_1 = -

0.029) in the high-volatility state, but the effects are not significantly different from that in the low volatility state.

Figure 2.7 shows the state-dependent impulse response function of the transitory component of stock prices to a one-standard deviation increase of Federal funds rate shock in Model 2. The response of the transitory component of stock prices to a monetary policy shock is larger when it is in the high-volatility state ($S_{et} = 1$). For example, a one positive standard deviation realization of monetary policy shock lowers stock price by a maximum amount of 0.78 when $S_{et} = 0$. However, when $S_{et} = 1$, the maximum response of stock price is much larger, reaching 1.88. The responses of stock prices to the monetary policy shock converge to zero in about three years.

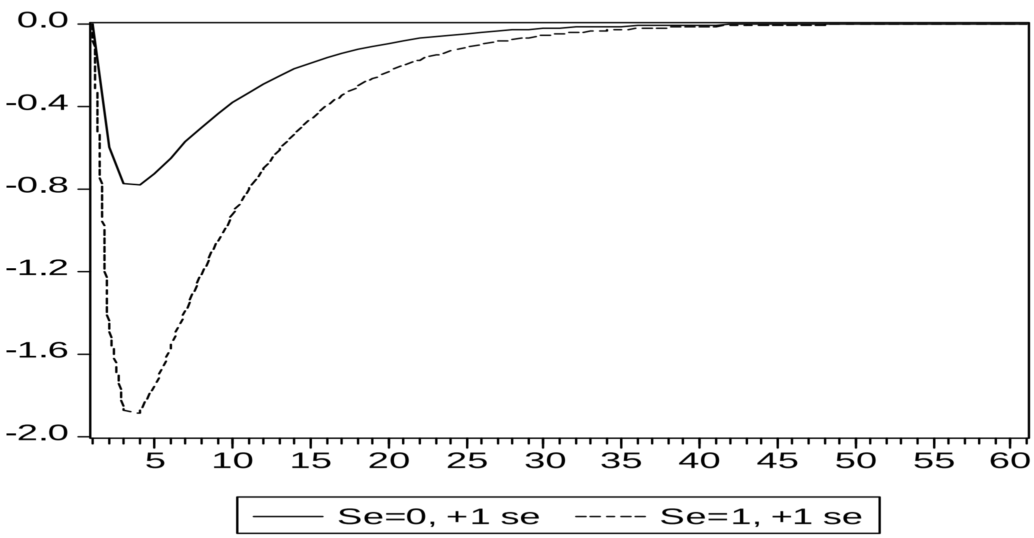

Figure 2.7. The state-dependent impulse response function of the transitory component of stock prices for small firms

2.4.3 The UC-MS Model with Monetary Policy and Time-Varying Transition Probabilities

In this section, I estimate the UC-MS model with monetary policy and also allowed the transition probabilities of the regime-switching process of S_{et} to be time-varying as equation (2.8) and (2.9), where time variation depends on monetary policy shocks. The results of estimation show that monetary policy has no significant effect on $P(S_{e,t} = 1 | S_{e,t-1} = 1)$, suggesting that the fixed transition probability is best for $P(S_{e,t} = 1 | S_{e,t-1} = 1)$. So, I estimate a model, named Model 3, which only $P(S_{e,t} = 0 | S_{e,t-1} = 0)$ is a function of Federal funds rate shocks. Results are showed in the fourth column in Table 2.1. The standard errors and transition probabilities of the trend component are all the same as Model 1 and Model 2. For the transitory component, a_0 is negative and significant, which implies that a contractionary monetary policy makes the transitory component of stock prices more likely to switch into the high volatility state. β_0 is also negative and significant. β_1 is a small negative number, but not significant. Thus, monetary policy has significant negative effects on stock prices in the low volatility state and the effects become a little larger in the high volatility state, however, the impact is no different between two states.

Figure 2.8a shows the stock prices and the estimated trend component of stock prices of Model 3. Figure 2.8b shows the estimated transitory component and the filtered probabilities of the trend component and the transitory component of Model 3. Since monetary policy can change the state of transitory component next period when it is in the low volatility in this period, the filtered probability of $S_e = 1$ becomes higher

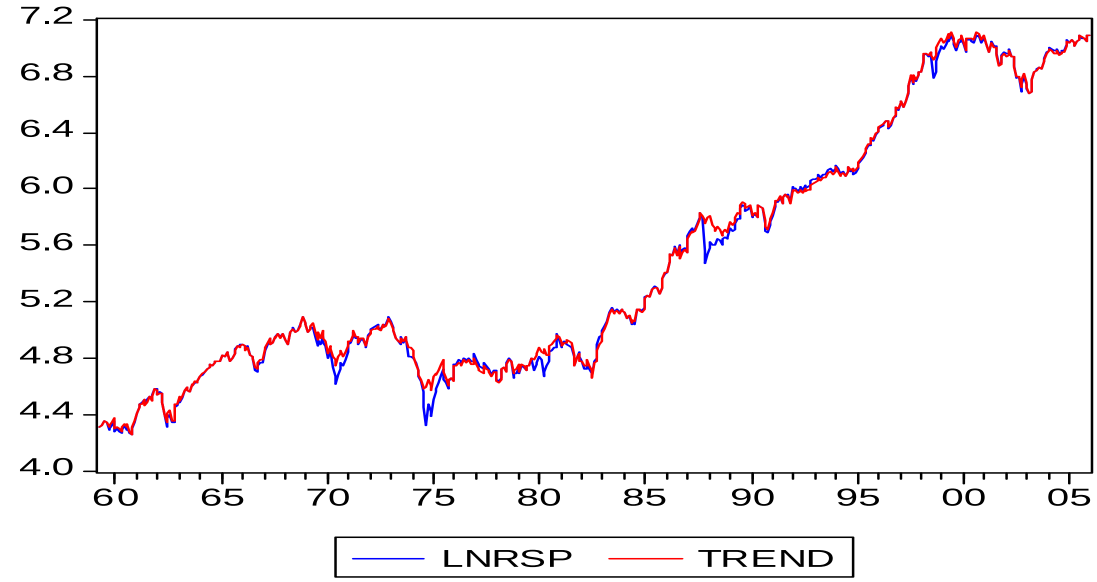

Figure 2.8a. Stock prices and the estimated trend component for Model 3

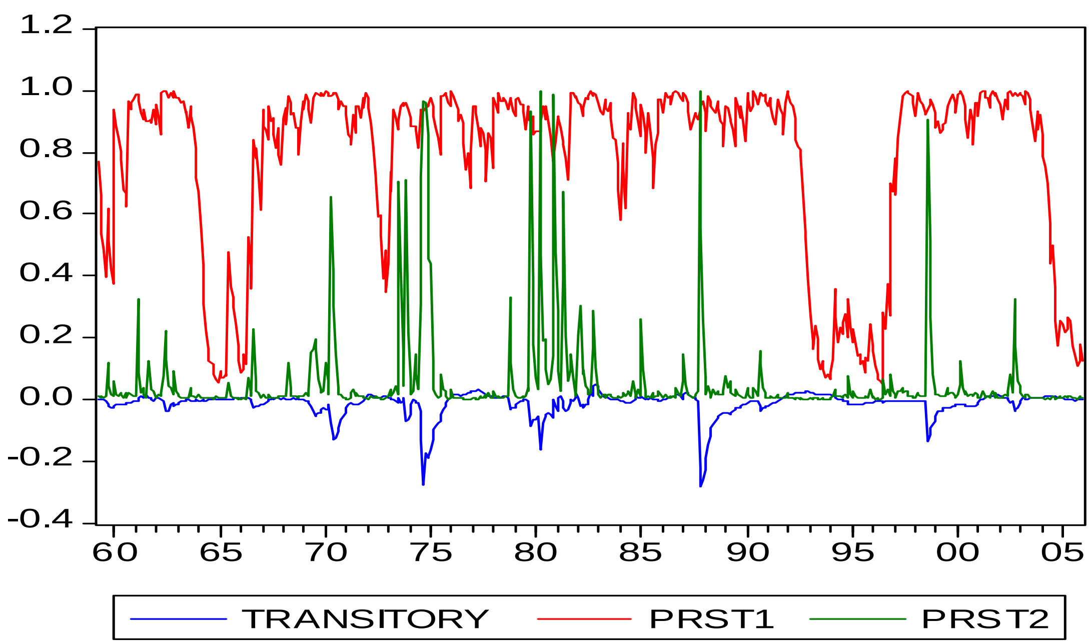

Figure 2.8b. The estimated transitory component and the filtered probabilities of the trend component and the transitory component for Model 2

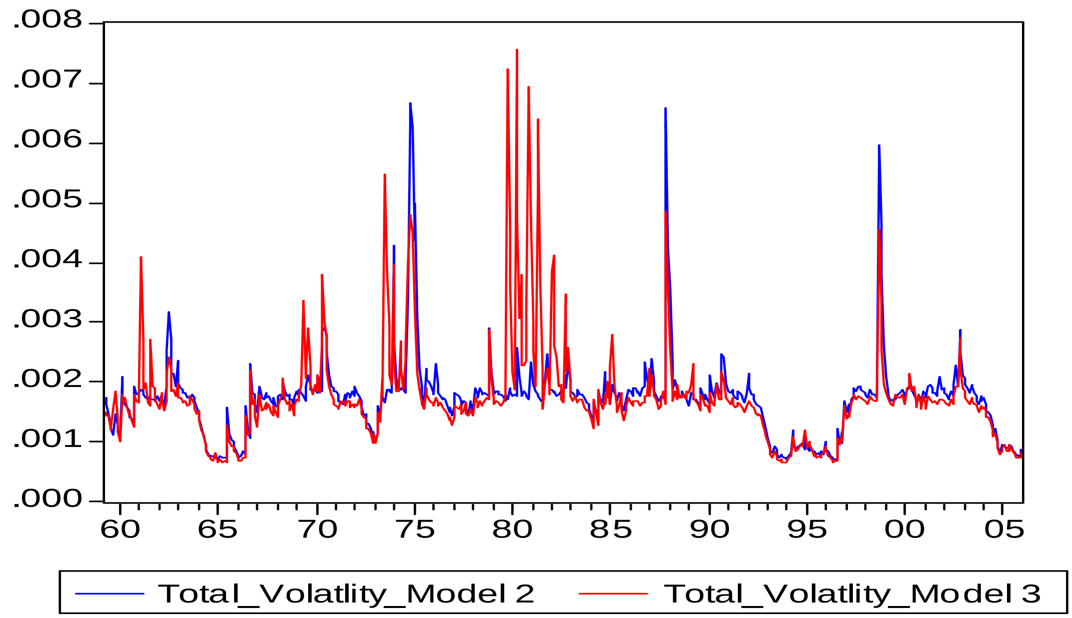

Figure 2.9a. Total volatility of Model 2 and Model 3

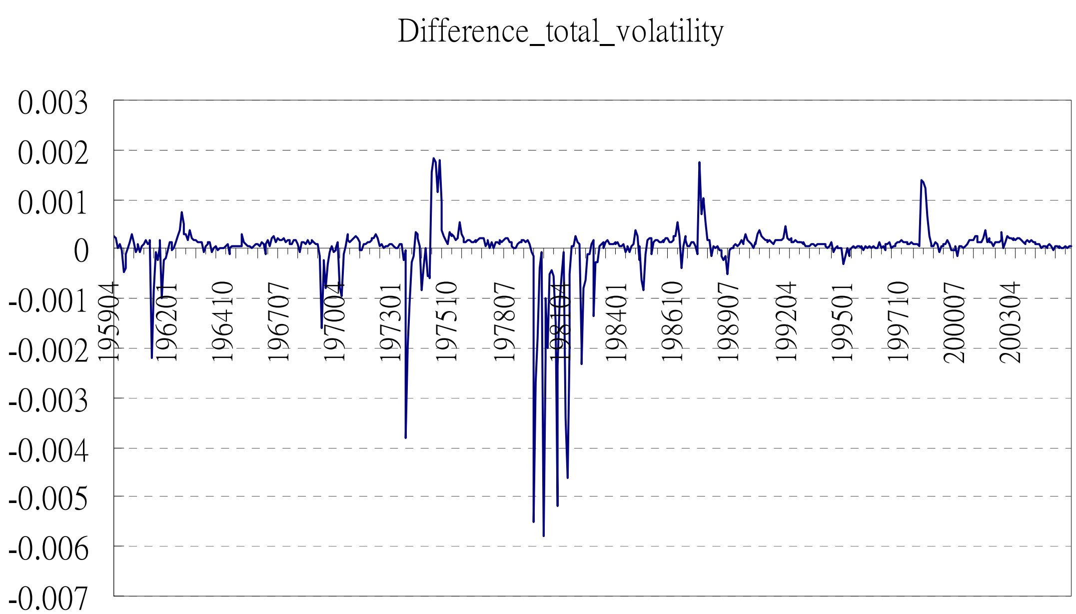

Figure 2.9b. Difference of total volatility between Model 2 and Model 3

comparing to Model 2, i.e. the probability of being in high volatility state at time t increases. It is especially obvious during October 1979 to November 1982.

Finally, I compare the volatility of Model 3 with Model 2. Figure 2.9a and Figure 2.9b show the total volatility of stock prices of Model 2 and Model 3 and the difference between these two volatilities, which is calculated as total volatility of stock prices of Model 2 minus that of Model 3. From Figure 2.10b, the difference of total volatility is normally around zero, but Model 3 increases the volatility during October 1979 to November 1982. Figure 2.10a, the graph of the volatility of the transitory component of Model 2 and Model 3, and Figure 2.10b, the graph of the difference between these two volatilities, show the same results. Therefore, Model 3 has no more explanatory ability to the volatility of stock market than Model 2.

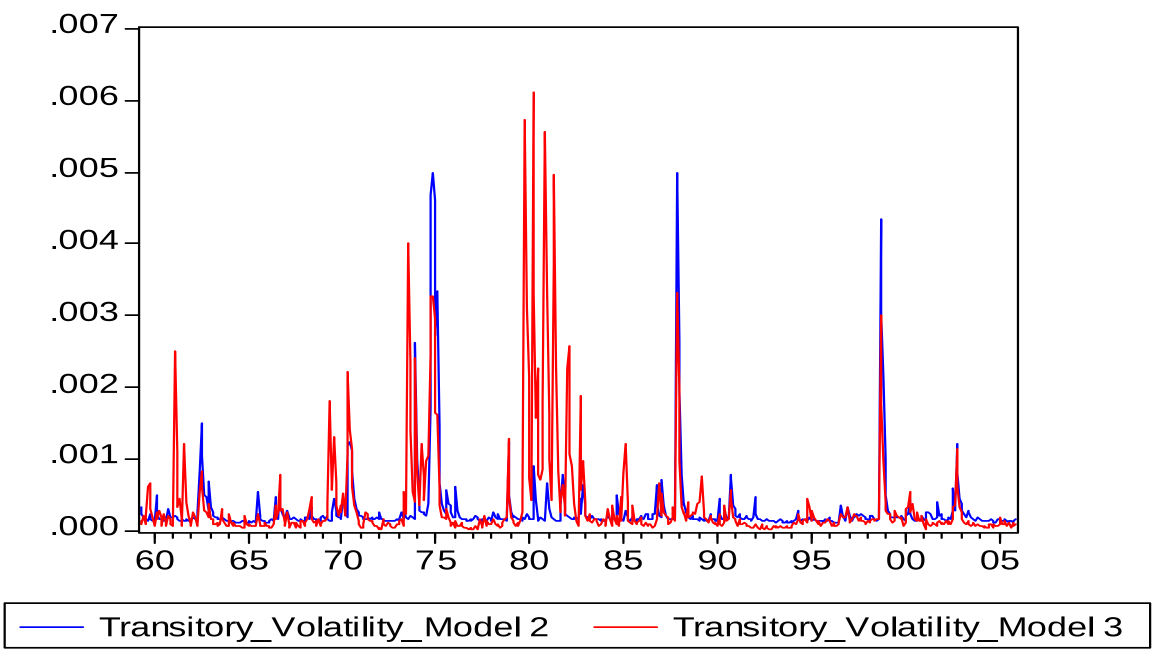

Figure 2.10a. Volatilities of the transitory component of Model 2 and Model 3

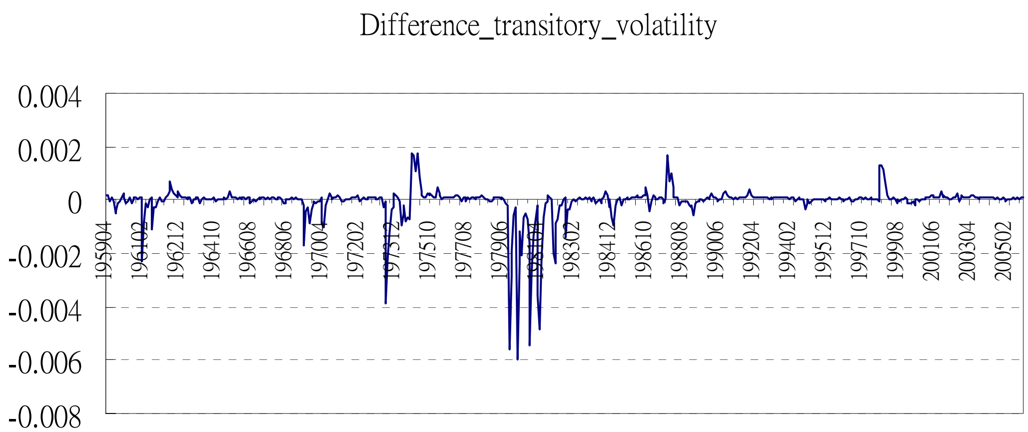

Figure 2.10b. Difference of volatility of the transitory component between Model 2 and Model 3

2.5 Conclusion

There is a considerable amount of papers investigating the effects of monetary policy on stock prices or stock returns, but there is still no consensus of conclusions about this question even though most of previous papers find the impact of monetary policy on stock prices is negative. From the theoretical point of view, a contractionary monetary policy can worse a firm's balance sheet positions and can reduce a firm's ability to borrow, spend and invest. This credit constraint problem is more likely to happen when the financial market is in a bad situation. This implies that monetary policy might have asymmetric effects on stock prices between different states of stock markets.

This paper attempts to investigate the asymmetric effects of monetary policy on stock market by using an unobserved-components model with Markov-switching heteroscedasticity. I decompose stock prices into the trend component and the transitory component and assume that monetary policy only influences the transitory component of stock prices. My results show that monetary policy has negative effects on stock prices, which is consistent with the most recent literature. When the transitory component is in the low volatility state, a contractionary monetary policy reduces stock prices and the effect is significant. When the transitory component is in the high volatility state, the negative effect of monetary policy becomes larger, but the difference of the monetary policy effects between two states is not significant. Besides, monetary policy can affect the dynamics of switching between low volatility and high volatility state. A contractionary monetary policy will lower the probability of staying in the low volatility state. Monetary policy also reduces the total volatility of stock prices and the volatility of the transitory component.

CHAPTER III

THE NONLINEAR EFFECT OF MONETARY POLICY ON STOCK RETURNS IN A SMOOTH TRANSITION AUTOREGRESSIVE MODEL

3.1 Introduction

The motivation of this chapter is the same as chapter II, but I try to investigate the nonlinear effect of monetary policy on stock market by using another empirical model. From pervious literature, not many paper focus on the question of whether monetary policy can have different effect on stock returns in different economic condition.

Hermann and Fratzscher (2004) present the evidence that the stock market response to monetary policy is highly asymmetric. They divide the 500 individual stocks comprising the S&P 500 into several groups according to the degree of financial constraints of firms and find the firms with more financial constraints are affected significantly more by monetary policy. Chen (2007) investigates the asymmetric monetary policy effects on stock returns by using Markov-switching models. He finds that monetary policy has larger effects on stock returns in bear market and a contractionary monetary policy leads to a higher probability of switching to the bear-market regime.

This paper attempts to investigate the nonlinear effect of monetary policy on stock returns by using the smooth transition autoregressive (STAR) model. Since output has a close relationship with stock returns and monetary policy and it might be the reason of nonlinear effects of monetary policy on stock returns, the growth rate of output is also included in our empirical model. The change in the Federal funds rate is used as an endogenous measure of monetary policy. Since stock returns, the change in the Federal funds rate, and the growth rate of output are all endogenously determined, the STAR models are constructed for all three variables.

One characteristic of this paper is that excess stock returns, the change in the Federal funds rate, and the growth rate of output are all allowed to be the possible threshold variable which controls for the nonlinear dynamics of models. Hence, this paper considers three asymmetries: asymmetry related to the state of stock market, asymmetry related to the direction and size of the monetary policy action, and asymmetry related to the state of economy. By appropriately choosing the best threshold variable for the model of each variable and estimating the nonlinear models for them, the nonlinear relationship among excess stock returns, monetary policy, and output growth can be investigated. Also, the nonlinear impulse response functions can help us to understand how they affect to each other.

My empirical results show that excess stock returns, monetary policy, and the growth rate of output all can be expressed in nonlinear STAR models. The threshold variable for the excess stock returns equation is the excess returns at lag two, the threshold variable for the change in the Federal funds rate equation is its own lag at two, and the threshold variable for the growth rate of output equation is also the excess stock returns at lag two, but with a different threshold value. The estimated coefficients and the impulse response functions show that the effect of monetary policy on excess returns of stock prices is significantly negative and nonlinear. The change in the Federal funds rate has a larger negative effect in the extreme low excess returns regime than in the high excess return regime. The possible explanation of this result might be that financial constraint of agents or firms are more likely to be bind when stock market is in a bad situation and excess returns are extremely low, so that monetary policy would have larger impact on stock returns in the low stock returns regime.

The rest of the chapter is organized as follows. Section 2 simply presents the framework of a STAR model that will be used in this paper and introduces the standard testing and estimating procedures. Section 3 is data and the empirical model. Section 4 reports the empirical results of estimating a set of nonlinear LSTAR models for excess returns, the change in the Federal funds rate, and the growth rate of output. Impulse response functions are reported in Section 5. Section 6 is the conclusion.

3.2 The STAR Model

3.2.1 The Basic Approach

The STAR model is a general form of the threshold autoregressive model which the transition variable is a function of the threshold variable, not just an indicator variable, so that the transition processes between regimes are smooth. A STAR model can be written as

$$y_t = a_0 + \alpha(L)x_t + [a_1 + \beta(L)x_t]F(z_{t-d}) + \varepsilon_t \quad (3.1)$$

where y_t is the dependent variable, x represents all the explanatory variables, including autoregressive lags of y_t, $F(z_{t-d})$ is the transition function, z_{t-d} is the transition variable that determines the switch between regimes, and d is the lag length of the transition variable. The dynamics of equation (1) changes with values of the transition variable. The nonlinear dynamics can be expressed as $\alpha(L)+\beta(L)F(z_{t-d})$.

Two common specifications for transition functions are the logistic and the exponential function. The logistic transition function is

$$F(z_{t-d}) = [1+e^{-\gamma(z_{t-d}-c)}]^{-1} \tag{3.2}$$

where γ determines the speed of transition and c is the threshold critical value. If $\gamma > 0$ ($\gamma < 0$), the logistic transition function changes smoothly from zero to one (from one to zero) when the transition variable z_{t-d} becomes increasingly larger than the threshold value c. The exponential STAR model has the transition function

$$F(z_{t-d}) = [1-e^{-\gamma(z_{t-d}-c)^2}], \quad \gamma > 0 \tag{3.3}$$

The exponential transition function smoothly approaches zero when the transition variable z_{t-d} is close to the threshold value c and approaches one when the transition variable z_{t-d} more deviates from the threshold value c.

3.2.2 Identifying and estimating methods

According to the STAR models developed by Luukkonen et al. (1988), Terasvirta and Anderson (1992), and Terasvirta (1994), there are four main steps to identify and estimate STAR models. First step is to identify and estimate a linear autoregressive model. The appropriate lag length for the model should be chosen before the tests of linearity and the estimation of the STAR models. In this paper, I search over different combinations of lags of explanatory variables and the multivariate lag length is selected based on the Schwarz Information Criteria (SIC). The multivariate lag length with the minimum SIC is chosen.

Second step is to identify possible candidates for the transition variable and test for the appropriateness of linearity. Terasvirta and Anderson (1992) propose an approximating equation and a procedure to test linear AR model against a nonlinear STAR model. The approximating equation of equation (1) can be expressed as:

$$y_t = c_0 + \phi_0(L)x_t + \phi_1(L)x_t z_{t-d} + \phi_2(L)x_t z_{t-d}^2 + \phi_3(L)x_t z_{t-d}^3 + v_t \tag{3.4}$$

The lag length of x_t was determined in the first step. For a given transition variable z and the amount of delay d, equation (4) can be estimated and also be tested for the hypothesis $\phi_1(L) = \phi_2(L) = \phi_3(L) = 0$. I repeat the estimation and hypothesis testing procedure for values of d from 1 to 4 in this paper. If there exists one or more values of d that reject the null hypothesis of linearity, it indicates a nonlinear STAR model and the delay d with the lowest probability value (i.e. the highest F-statistic) is chosen.

The third step is to identify the specification of STAR model. If the null hypothesis of linearity is rejected and the transition variable is determined, the specification of STAR model must be chosen between logistic STAR and exponential STAR model. A sequence of hypothesis tests and decision rules based on equation (4) proposed by Terasvirta and Anderson (1992) are:

$$\mathrm{H}_{0,1}: \phi_3(L) = 0 \tag{3.5}$$

$$\mathrm{H}_{0,2}: \phi_2(L) = 0 \mid \phi_3(L) = 0 \tag{3.6}$$

$$\mathrm{H}_{0,3}: \phi_1(L) = 0 \mid \phi_2(L) = \phi_3(L) = 0 \tag{3.7}$$

If $\mathrm{H}_{0,1}$ is rejected, select an LSTAR model. If $\mathrm{H}_{0,1}$ is not rejected and $\mathrm{H}_{0,2}$ is rejected, select an ESTAR model. If $\mathrm{H}_{0,1}$ and $\mathrm{H}_{0,2}$ are not rejected but $\mathrm{H}_{0,3}$ is rejected, select an LSTAR model.

The final step is the estimation of the STAR model. The threshold value c and the rate of transition between regimes γ are determined by a two-dimensional grid search over data points of the transition variable z and different values of γ. The combination values of z and γ with the minimum sum of squared errors are the optimal estimates. Then, the model can be estimated by using nonlinear least squares.

3.3 Data and the Empirical Model

3.3.1 Data

Monthly data on the Standard and Poors 500 stock index is used as the stock prices. The excess returns of stock prices (XR) are defined as the monthly percentage change in the S&P 500 index minus the monthly yield on 3-month U.S. Treasury Bill. The change in the Federal funds rate (DFF) is used as the monetary policy variable and is calculated as the first difference of the Federal funds rate which is already divided by 12 for the monthly frequency. Since output has a close relationship with stock returns and monetary policy and it might be the reason of nonlinear effects of monetary policy on stock returns, the growth rate of output is also included in our empirical model. Industrial production is used as output for the monthly purpose, while real Gross Domestic Product is only available at quarterly frequency. The growth rate of output is measured by the percentage change in industrial production.

All original data are collected from Federal Reserve Economic Data and the Center for Research in Security Prices (CRSP) database and start from July 1954 to December 2005. Due to the lost of one data points in calculating the change in the Federal funds rate, the values of excess returns, the change in the Federal funds rate, and the growth rate of industrial production are from August 1954 to December 2005 and all expressed in percentage terms. Figure 3.1 shows the time series plot of excess stock returns (XR), the change in the Federal funds rate (DFF), and the growth rate of output (Gy) from August 1954 to December 2005. Figure 3.2 shows the scatter plot of excess stock returns and the change in the Federal funds rate. From Figure 1 the time series plot of the change in the Federal funds rate, it shows a very different pattern during the period of October 1979 to October 1982, especially the fluctuation of DFF in this period is larger than the rest of the sample period. Therefore, I will include a dummy variable for this special period in the model.

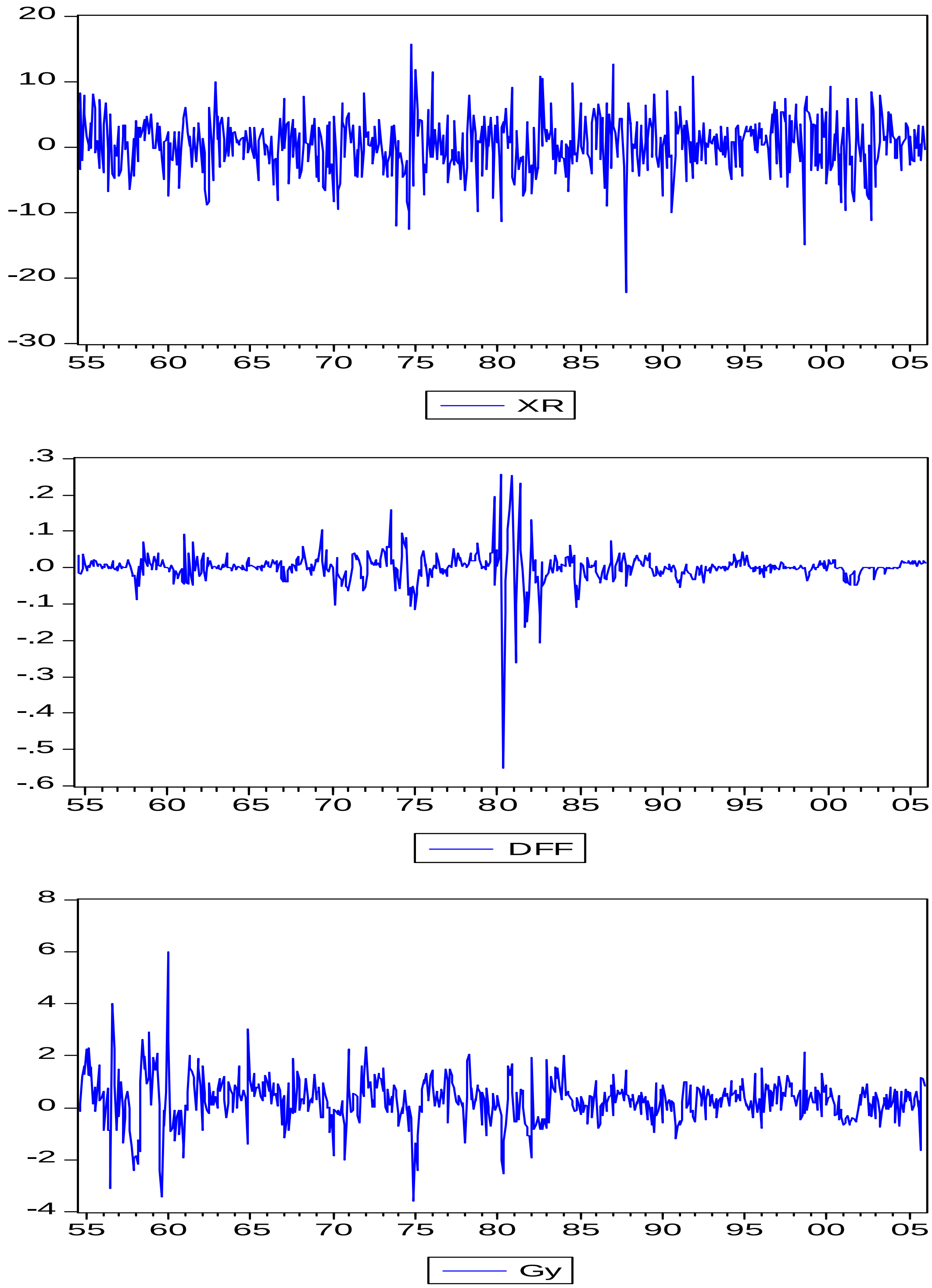

Figure 3.1. The time series plot of excess stock returns, the change in the Federal funds rate, and the growth rate of industrial production: 1954:8 – 2005:12

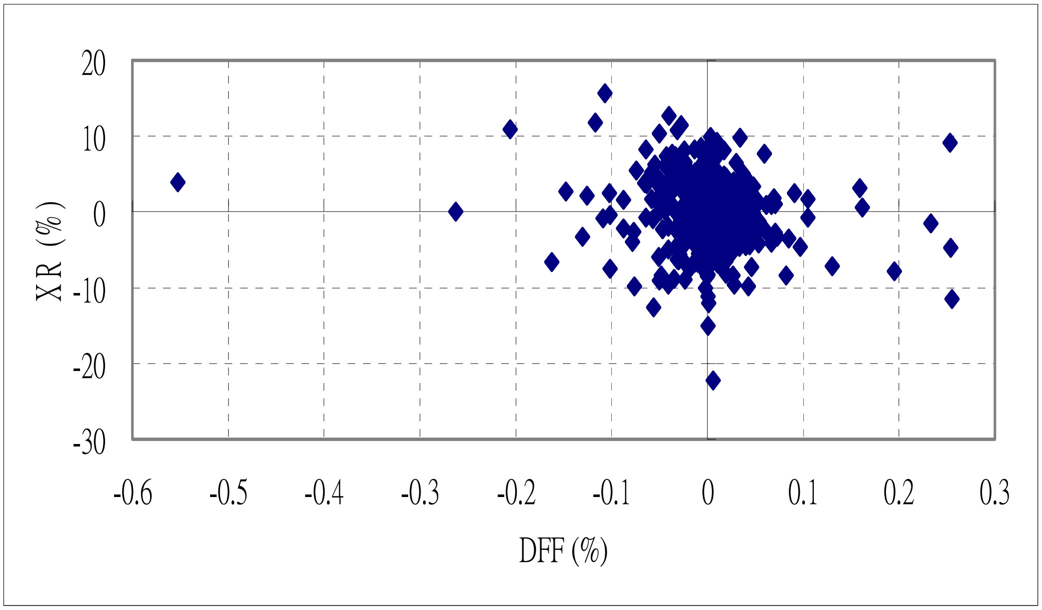

Figure 3.2. The scatter plot of the scatter plot of excess stock returns and the change in the Federal funds rate: 1954:8 ~ 2005:12

The other thing is that the monthly data of stock returns downloaded from CRSP database are values on the last day of each month, but the data of Federal funds rate and industrial production are averages of daily figures of a month. When we look at the time path in Figure 3.3, the Federal funds rate and industrial production represent the values in the middle of a month and excess stock returns are values in the end of a month. Considering the quick reaction of the stock market, I allow the change in the Federal funds rate and the growth rate of industrial production can have contemporaneous effects on the excess stock returns in the model.

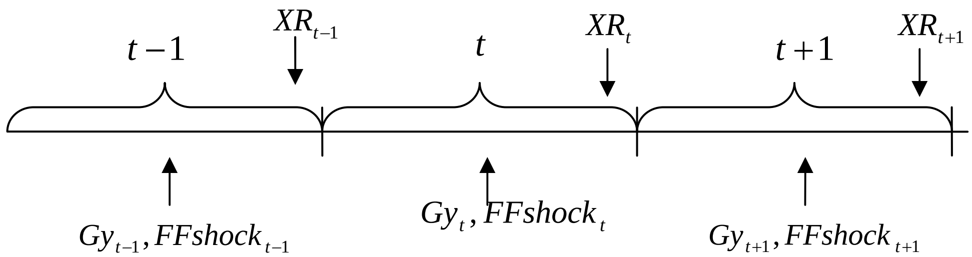

Figure 3.3. Time path of variables

3.3.2 The empirical model

Since excess stock returns, the change in the Federal funds rate, and the growth rate of output are all endogenously determined, STAR models for these three variables can be written as:

$$XR_t = a_0 + \sum_{i=1}^{p_1} \alpha_{1i} XR_{t-i} + \sum_{i=0}^{q_1} \alpha_{2i} DFF_{t-i} + \sum_{i=0}^{q_1} \alpha_{3i} DFF_{t-i} * Dummy_{t-i} + \sum_{i=0}^{r_1} \alpha_{4i} Gy_{t-i}$$

$$+ [a_1 + \sum_{i=1}^{p_1} \alpha_{5i} XR_{t-i} + \sum_{i=0}^{q_1} \alpha_{6i} DFF_{t-i} + \sum_{i=0}^{q_1} \alpha_{7i} DFF_{t-i} * Dummy_{t-i}$$

$$+ \sum_{i=0}^{r_1} \alpha_{8i} Gy_{t-i}] F(z_{1,t-d}) + \varepsilon_{1t} \tag{3.8}$$

$$DFF_t = b_0 + \sum_{i=1}^{p_2} \beta_{1i} XR_{t-i} + \sum_{i=1}^{q_2} \beta_{2i} DFF_{t-i} + \sum_{i=1}^{r_2} \beta_{3i} Gy_{t-i}$$

$$+ [b_1 + \sum_{i=1}^{p_2} \beta_{4i} XR_{t-i} + \sum_{i=0}^{q_2} \beta_{5i} DFF_{t-i} + \sum_{i=0}^{r_2} \beta_{6i} Gy_{t-i}] F(z_{2,t-d}) + \varepsilon_{2t} \tag{3.9}$$

$$Gy_t = c_0 + \sum_{i=1}^{p_3} \gamma_{1i} XR_{t-i} + \sum_{i=1}^{q_3} \gamma_{2i} DFF_{t-i} + \sum_{i=1}^{q_3} \gamma_{3i} DFF_{t-i} * Dummy_{t-i} + \sum_{i=1}^{r_3} \gamma_{4i} Gy_{t-i}$$

$$+[c_1+\sum_{i=1}^{p_3}\gamma_{5i}XR_{t-i}+\sum_{i=1}^{q_3}\gamma_{6i}DFF_{t-i}+\sum_{i=1}^{q_3}\gamma_{7i}DFF_{t-i}*Dummy_{t-i}$$

$$+\sum_{i=1}^{r_3}\gamma_{8i}Gy_{t-i}]F(z_{3,t-d})+\varepsilon_{3t} \tag{3.10}$$

where XR_t is excess returns of stock prices, DFF_t is the change in the Federal funds rate, Gy_t is the growth rate of industrial production, $Dummy_t$ is a time dummy variable for the special period October 1979 to October 1982: $Dummy_t$ = 1, if t = October 1979 to October 1982, named the special period; 0, if t = August 1954 to September 1979 and November 1982 to December 2005, named normal period. p_i, q_i, r_i, for i =1, 2, 3 are lag length of XR_t, DFF_t, and Gy_t in equation (3.8), (3.9), and (3.10), respectively. z_{1t-d}, z_{2t-d}, and z_{3t-d} are transition variables for XR_t, DFF_t, and Gy_t equations, respectively. XR_{t-d}, Gy_{t-d}, and DFF_{t-d} are all considered as the possible transition variable for each equation. The lag length of the transition variable d is chosen from 1 to 4. Transition functions $F(z_{1t-d})$, $F(z_{2t-d})$, and $F(z_{2t-d})$ are all selected between the logistic function and the exponential function.

For the model of excess stock returns in equation (3.8), the change in the Federal funds rate and the growth rate of output are allowed to have contemporaneous effects on the excess stock returns, so DFF_t, and Gy_t are included in explanatory variables. Moreover, to capture the possible different effects of monetary policy DFF_t on excess stock returns between the special period and the normal period, a dummy variable $Dummy_t$ is included in conjunction with DFF_t. Thus, the effect of DFF_{t-i} on XR_t is nonlinear: it varies between α_{2i} and $\alpha_{2i} + \alpha_{6i}$ in the normal period and varies between $\alpha_{2i} + \alpha_{3i}$ and $\alpha_{2i} + \alpha_{3i} + \alpha_{6i} + \alpha_{7i}$ in the special period.

For the model of Federal funds rate in equation (3.9), the change in the Federal funds rate DFF_t is a function of lags of excess stock returns XR_{t-i}, lags of the change in the Federal funds rate DFF_{t-i}, lags of growth rate of industrial production Gy_{t-i}, and the transition variable z_{2t-d}. For dealing with the special period for DFF_t, if coefficients of all explanatory variables are allowed to change between two sample periods, there might be too many regressors in the equation. Therefore, I adopt the generalized least squares estimation method to deal with the high volatility of DFF_t in the special period. I first estimate the linear AR model of equation (3.9) and get the residuals. I then calculate the standard deviation of the residuals for the special period and for the normal period, so that the ratio of the standard deviation of the residuals of the special period to the normal period can be known. Next, divide each observation (both dependent and explanatory variables) in the special period by the ratio in order to shrink the high volatility in that period. Finally, the four steps of identifying and estimating STAR models can be applied to the transformed dependent and explanatory variables.

For the model of industrial production in equation (3.10), the explanatory variables of the growth rate of industrial production Gy_t are lags of excess stock returns XR_{t-i}, lags of the change in the Federal funds rate DFF_{t-i}, lags of growth rate of industrial production Gy_{t-i}, and the transition variable z_{3t-d}. As the same season in the excess returns equation, the interaction term $DFF_{t-i} * Dummy_{t-i}$ is included in the model as well.

3.4 Empirical Results

Following the four main steps of identifying and estimating STAR models that have mentioned in Section 2.2, the empirical results of STAR models for the excess returns of stock prices, the change in the Federal funds rate, and the growth rate of industrial production are discussed as follows.

First step is choosing the appropriate lag length in order to identify and estimate linear autoregressive models for each series. The multivariate lag length is selected based on the Schwarz Information Criteria (SIC) and the lag length with the minimum SIC is chosen. For the excess stock returns equation, the SIC indicates a model with one autoregressive lag of excess returns, the change in the Federal funds rate at time t, and the growth rate of industrial production at time t. For the Federal funds rate equation, the SIC indicates a model with two lags of excess return, two lags of the change in the Federal funds rate, and one lag of industrial production growth. For the growth rate of industrial production equation, two lags of excess returns, one lag of the change in the Federal funds rate, and one lag of industrial production growth are chosen to be the explanatory variables.

Second step is to test for the appropriateness of linearity. Lags of XR_{t-d}, DFF_t, and Gy_t are all considered as candidates for the transition variable, besides, DFF_t and Gy_t are allowed to be the transition variable for excess stock returns equation. Results of linearity test are showed in Table 3.1a for the excess stock returns equation, Table 3.1b for the Federal funds rate change equation, and Table 3.1c for the growth rate of output equation.

Table 3.1a Linearity test and determination of lag order for transition variable — XR equation

Transition variabl	XR_{t-d}		DFF_{t-d}		Gy_{t-d}	
	F	p-value	F	p-value	F	p-value
d =0			1.6024	0.0865	0.6718	0.7793
d =1	1.1068	0.3514	0.8279	0.6216	1.8476	0.0381
d =2	**2.3203**	**0.0067**	2.2842	0.0077	1.1778	0.2953
d =3	1.8217	0.0416	0.8547	0.5937	0.8376	0.6114
d = 4	1.2342	0.2553	1.6804	0.0671	0.7989	0.6518

Table 3.1b Linearity test and determination of lag order for transition variable — DFF equation

Transition variab	XR_{t-d}		DFF_{t-d}		Gy_{t-d}	
	F	p-value	F	p-value	F	p-value
d =1	1.6221	0.0633	4.2478	0.0000	2.4850	0.0015
d =2	3.4073	0.0000	2.3484	0.0028	2.5865	0.0009
d =3	2.0227	0.0122	**4.3173**	**0.0000**	1.8072	0.0304
d = 4	1.8866	0.0219	2.3321	0.0030	1.6591	0.0549

Table 3.1c Linearity test and determination of lag order for transition variable —Gy equation

Transition variabl	XR_{t-d}		DFF_{t-d}		Gy_{t-d}	
	F	p-value	F	p-value	F	p-value
d =1	2.0825	0.0094	2.8401	0.0003	3.3473	0.0000
d =2	**4.1848**	**0.0000**	3.1811	0.0000	1.5439	0.0848
d =3	2.0449	0.0111	1.6860	0.0494	2.3996	0.0022
d = 4	1.6123	0.0657	1.3110	0.1896	2.9341	0.0002

For the excess stock returns equation, Table 3.1a shows that when XR_{t-d} is the transition variable, d = 2 and d =3 both reject the null hypothesis of linearity at a 5%

significant level with p-value 0.0067 and 0.0416, respectively; when DFF_{t-d} is the transition variable, d = 2 rejects the linearity with p-value 0.0077; When Gy_{t-d} is the transition variable, d = 1 rejects the linearity with p-value 0.0381. Thus, XR_{t-2} that gives the lowest marginal probability value (highest F-statistics) is chosen to be the transition variable in the excess stock returns equation. By using the same choosing criterion, Table 3.1b shows that DFF_{t-3} that gives the highest F-statistics is the best transition variable for the Federal funds rate equation. Table 3.1c shows that XR_{t-2} is the best transition variable for the growth rate of industrial production.

Table 3.2 Tests for the STAR specification

The excess stock returns equation with the transition variable XR_{t-2}

	F-statistic	p-value
$\mathrm{H}_{0,1}: \phi_3(L) = 0$	2.6723	0.0313

The change in the Federal funds rate equation with the transition variable DFF_{t-2}

	F-statistic	p-value
$\mathrm{H}_{0,1}: \phi_3(L) = 0$	3.2611	0.0065

The growth in industrial production equation with the transition variable XR_{t-2}

	F-statistic	p-value
$\mathrm{H}_{0,1}: \phi_3(L) = 0$	2.0916	0.0648
$\mathrm{H}_{0,1}: \phi_2(L) = 0 \mid \phi_3(L) = 0$	6.7142	0.0000

Since all three variables are accepted to be modeled as STAR models, the next step is to identify the specification of STAR model. A sequence of hypothesis testing in equation (3.5) - (3.7) based on equation (3.4) is tested and the results are reported in Table 3.2. The results of hypothesis tests for the STAR model specification in the excess stock returns equation are reported in Table 3.2. For excess returns equation and the Federal funds rate equation, the hypothesis $H_{0,1}$ is significantly rejected at a 5% significant level, which indicates the choice of LSTAR models for both variables. For the growth rate of industrial production equation, the hypothesis $H_{0,1}: \phi_3(L) = 0$ is not rejected at a 5% significant level, but $H_{0,2}: \phi_2(L) = 0 \mid \phi_3(L) = 0$ is significantly rejected. So, according to the decision rules, it indicates that an ESTAR model is the appropriate specification for the growth rate of industrial production. Then, let's turn to the estimation of models.

Table 3.3 Model estimates of excess returns of stock prices

	Linear	LSTAR (XR_{t-2})
a_0	0.224 (0.177)	1.312 (1.321)
XR_{t-1}	0.013 (0.040)	0.051 (0.161)
DFF_t	-22.101 (6.569)**	-98.613 (39.530)**
$DFF_t * Dummy_t$	12.074 (7.806)	75.467 (36.363)**
Gy_t	0.145 (0.197)	4.029 (1.368)**
$a_1 * F(XR_{t-2})$		-1.119 (1.334)
$XR_{t-1} * F(XR_{t-2})$		-0.010 (0.167)
$DFF_t * F(XR_{t-2})$		80.898 (40.236)**
$DFF_t * Dummy_t * F(XR_{t-2})$		-69.008 (37.490)*
$Gy_t * F(XR_{t-2})$		-3.981 (1.382)**
γ		23.15
c		-8.4
R^2	0.026	0.049
Log likelihood	-1751.455	-1741.270

Note: Standard errors are in the parentheses. Asterisk (*) denotes statistically significant at the 10% level. Double-asterisk (**) denotes statistically significant at the 5% level.

The results of model estimates of excess stock returns are shown in Table 3.3. The second column reports the estimates of the linear AR model. It indicates that the constant term, lagged excess returns, and the growth rate of output have no significant explanatory ability to excess returns, but the change in the Federal funds rate has a negative and statistically significant effect on excess returns. A contractionary monetary policy reduces excess stock returns, which is consistent with the conclusions of most of the recent literature. The coefficient of the interaction term $DFF_t * Dummy_t$ is positive but not significant, which means that the monetary policy is not different between the special period and the normal period.

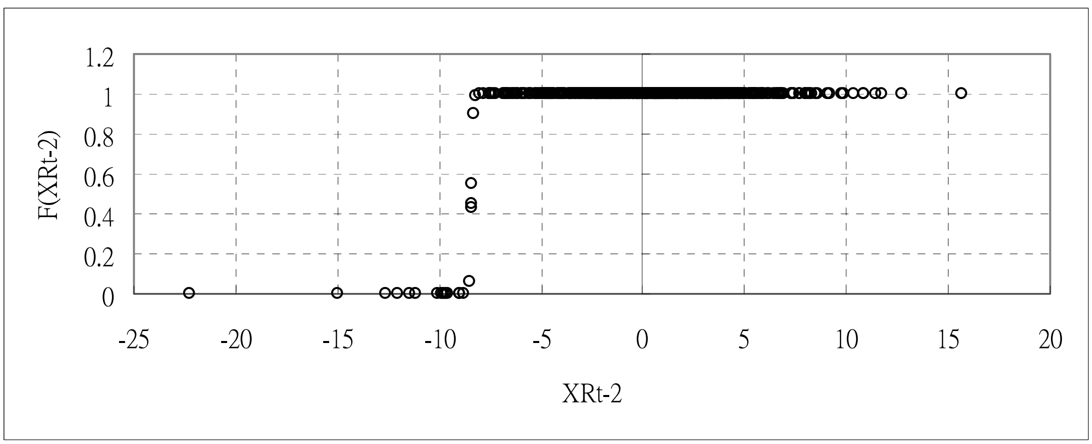

Figure 3.4. The logistic transition function for XR equation: $F = [1 + e^{-23.15(XR_{t-2}-(-8.4))}]^{-1}$

The third column in Table 3.3 shows estimates of excess returns in a nonlinear LSTAR model when the transition variable is the excess stock returns at lag two. The estimated transition value is -8.4% of excess return of stock prices. The rate of transition γ is 23.15, which indicates that the model switches between regimes very fast. Figure

3.4 shows the plot of the estimated transition function of excess returns. Since the rate of transition is large, the equation of excess stock returns seems to be explained in two different regimes. The majority of data points are in the high excess return regime which XR is larger than transition value -8.4%, and only a few data points are in the low excess returns regime which XR is smaller than transition value -8.4%.

When excess stock returns are modeled in the LSTAR model, monetary policy and the growth rate of industrial production have significant explanatory ability. In the low excess stock returns regime, the change in the Federal funds rate DFF_t has large negative impact on excess returns. If the change in the Federal funds rate increases 1% (i.e. 12% increases in annual rate), the excess stock returns will decrease 98.613%. In the high excess returns regime, the coefficient of DFF_t is -98.613 + 80.989 = -17.624, the effects of the change in the Federal funds rate on excess returns becomes smaller. So, we can know that the monetary policy has larger effects on excess returns when excess returns are very low. The possible explanation of this result might be that financial constraint of agents are more likely to be bind when stock market is bad and excess returns are extremely low, so that monetary policy would have larger impact on stock returns in the extreme low stock returns regime. From the coefficient of $DFF_t * Dummy_t$, the response of excess returns to monetary policy is smaller in the special period, October 1979 to October 1982, than in the normal period, but still the response is larger in the extreme lower excess returns regime than in the high excess returns regime.

The effects of the growth rate of industrial production on excess stock returns are also nonlinear. The coefficient of the growth rate is 4.029 in the extreme low excess returns regime, but is 4.029 - 3.981 = 0.048 in the usual excess returns regime. This shows that the when the excess return is very low, an increase in the growth rate of industrial production has can increase the excess returns a lot, but when the excess return is larger than -8.4%, an increase in the growth rate of industrial production does not have a big impact on excess returns.

Table 3.4 Model estimates of the change in Federal funds rate

	Linear	LSTAR ($z = DFF'_{t-3}$)
a_0	-0.0012 (0.0010)	0.0060 (0.0062)
XR_{t-1}	0.00008 (0.00024)	-0.0007 (0.0012)
XR_{t-2}	0.0006 (0.0002)**	0.0008 (0.0009)
DFF_{t-1}	0.3489 (0.0403)**	0.3571 (0.1653)**
DFF_{t-2}	0.1031 (0.0400)**	-0.0723 (0.1942)
Gy_{t-1}	0.0049 (0.0012)**	0.0255 (0.0050)**
$a_1 * F(DFF'_{t-3})$		-0.0067 (0.0063)
$XR_{t-1} * F(DFF'_{t-3})$		0.0007 (0.0012)
$XR_{t-2} * F(DFF'_{t-3})$		-0.0002 (0.0009)
$DFF_{t-1} * F(DFF'_{t-3})$		-0.0168 (0.1715)
$DFF_{t-2} * F(DFF'_{t-3})$		0.1696 (0.2006)
$Gy_{t-1} * F(DFF'_{t-3})$		-0.0221 (0.0052)**
γ		846
c		-0.048
R^2	0.219	0.248
Log likelihood	1420.879	1430.099

Note: Standard errors are in the parentheses. Asterisk (*) denotes statistically significant at the 10% level. Double-asterisk (**) denotes statistically significant at the 5% level.

Table 3.4 reports the estimates of models of the change in the Federal funds rate. The estimates of the linear model are reported in the second column. The coefficients of XR_{t-1} and XR_{t-2} are positive, which indicates that when the lags of excess returns increase, monetary authority will adopt a contractionary monetary policy. The coefficients of Gy_{t-1} is significant positive, so the Federal funds rate also reacts positively to the growth rate of industrial production. Those are consistent with the expectation.

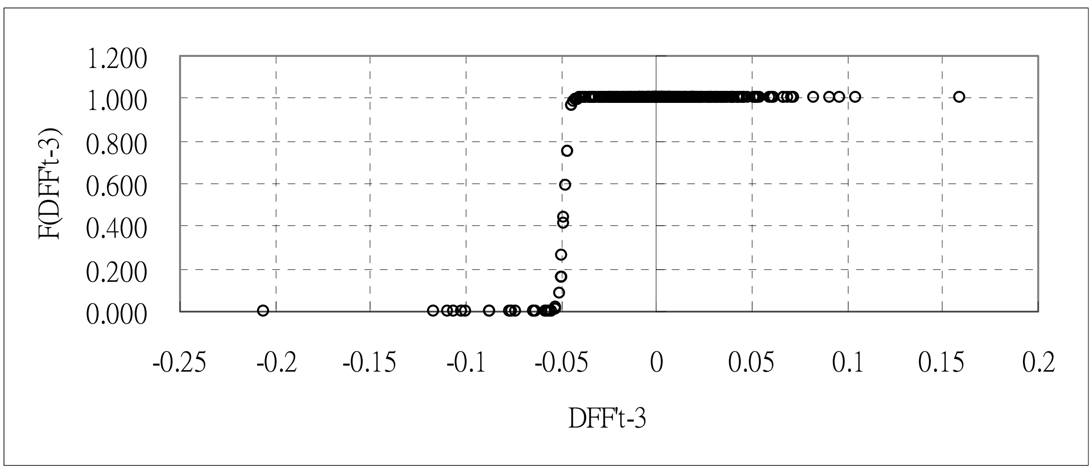

Figure 3.5. The logistic transition function for DFF equation: $F = [1 + e^{-846(DFF'_{t-3}-(-0.048))}]^{-1}$

The third column in Table 3.4 presents estimates of the LSTAR model with the transition variable DFF_{t-3} for the change in the Federal funds rate. The estimated transition value is -0.048 and the rate of transition γ is 846. Figure 3.5 shows the graph of the transition function for the Federal funds rate equation. The model shifts between regimes quickly because of the large value of rate of transition. Most of the observations are in the regime of $DFF_{t-3} > -0.048$, called usual monetary policy regime. Only few observations in the regime of $DFF_{t-3} \leq -0.048$, named the extreme expansionary monetary policy regime.

Lagged excess stock returns have no significant effects on monetary policy actions. Monetary authority only significantly responds to DFF_{t-1} and Gy_{t-1}. The effect of growth rate of output on monetary policy is similar. The coefficient of Gy_{t-1} is 0.0255 when $F(DFF_{t-3}) = 0$ and is 0.0034 when $F(DFF_{t-3}) = 1$. The positive response of the change in the Federal funds rate to the growth rate of industrial production is larger in the extreme expansionary monetary policy regime. In the usual monetary policy regime, the reaction of the monetary policy to the growth rate of industrial production is almost zero. The possible explanation for the results is that when there was an extreme expansionary monetary policy in three periods ago (the transition variable DFF_{t-3} is in lower regime), money supply increases and induces inflation and economy, therefore, monetary authority will become more sensitive to the change of its target, including the change in the growth rate of output.

Table 3.5 presents the model estimates of the growth rate of output. The estimates of the linear model in second column show that all explanatory variables are positive related to the growth rate of output, but the estimates of an ESTAR model in the third column show a different result. The estimated transition value of XR_{t-2} is 3.1% and the rate of transition γ is 0.008. Figure 3.6 shows the graph of the transition function for the growth rate of output. Most of the observations are in the regime that

$F(XR_{t-2})$ is less than 0.5, named the normal excess returns regime. Only a few observations are in the extreme excess returns regime (extreme low or extreme high).

When excess returns at lag two is close to the threshold value 3.1%, the growth rate of industrial production responds positively to the excess returns. This is for the majority of the data. But when excess returns at lag two deviate from the threshold value 3.1%, the relationship between lagged excess stock return and the growth rate of industrial production becomes negative. The real reason of how lagged excess stock return help explain the growth rate of output is not sure. The possible explanation is that

Table 3.5 Model estimates of growth in industrial production

	Linear	ESTAR (XR_{t-2})
a_0	0.168 (0.035)**	0.305 (0.055)**
XR_{t-1}	0.015 (0.008)*	0.032 (0.010)**
XR_{t-2}	0.023 (0.008)**	0.019 (0.014)
DFF_{t-1}	3.601 (1.294)**	1.696 (1.643)
$DFF_{t-1} * Dummy_{t-1}$	-3.124 (1.527)**	0.611 (1.938)
Gy_{t-1}	0.352 (0.038)**	0.210 (0.047)**
$a_1 * F(XR_{t-2})$		-0.958 (0.279)**
$XR_{t-1} * F(XR_{t-2})$		-0.085 (0.033)**
$XR_{t-2} * F(XR_{t-2})$		-0.057 (0.031)*
$DFF_{t-1} * F(XR_{t-2})$		6.382 (6.170)
$DFF_{t-1} * Dummy_{t-1} * F(XR_{t-2})$		-16.708 (7.853)**
$Gy_{t-1} * F(XR_{t-2})$		0.852 (0.210)**
γ		0.0075
c		3.1
R^2	0.172	0.235
Log likelihood	-744.561	-720.171

Note: Standard errors are in the parentheses. Asterisk (*) denotes statistically significant at the 10% level. Double-asterisk (**) denotes statistically significant at the 5% level.

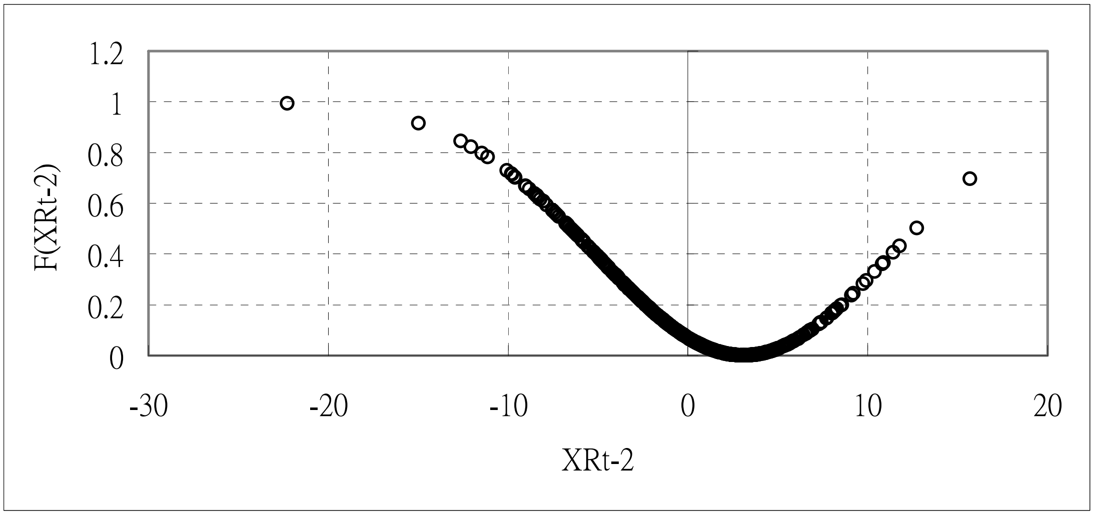

Figure 3.6. The exponential transition function for Gy equation: $F = [1 - e^{-0.008\,(XR_{t-2}-3.1)^2}]$

stock returns might reflect quickly to the information about the future movements of output, so normally stock returns are positively related to the growth rate of output. When XR_{t-2} is extremely high or low, the excess stock returns of lag one might have a negative relationship with the growth rate of output, but this is only for a few extreme excess stock returns in the sample.

From the estimates of DFF_{t-1}, monetary policy has positive but not significant effects on the growth rate of industrial production in the normal period. But, in the special period, the effect of monetary policy on the growth rate of industrial production is negative (1.696 + 0.611 + 6.382 - 16.708 = -8.019) and significant with p-value of 0.075 in Wald test when the excess stock returns are in the extreme regime. This implies that the monetary policy is not neutral to the growth rate of industrial production only in the extreme excess stock returns regime for the special period.

3.5 Impulse Response Functions

In order to understand how excess returns of stock prices, the growth rate of output, and the change in the Federal funds rate respond to shocks, nonlinear impulse response functions for the estimated LSTAR models are calculated. Before calculating the impulse response functions, the assumptions to solve the identification problem in our three equations model need to be discussed first. The monthly data of stock returns are values on the last day of each month, but the data of Federal funds rate and industrial production are averages of daily figures of a month, so I assume that the change in the Federal funds rate and the growth rate of industrial production can have contemporaneous effects on the excess stock returns, but not vice versa. Since the change in the Federal funds rate at time t and the growth rate of industrial production at time t are already included in the excess returns equation, the error term of the excess returns equation ε_{1t} is also the real excess returns disturbance.

The additional identification restriction is to use Choleski decomposition for the change in the Federal funds rate and the growth rate of industrial production. I constrain the system such that the contemporaneous value of the change in Federal funds rate does not have a contemporaneous effect on the growth rate of industrial production. The order of output place in front of monetary policy action is used in most literature. Thus, the impulse response functions in the system are calculated by using the following order: (Gy_t, DFF_t, XR_t). It can also be explained by relationship between the error terms of the equation (3.8) to (3.10) and the underlying economic shocks u_{it}: $\varepsilon_{1t} = u_{1t}$, $\varepsilon_{2t} = bu_{3t} + u_{2t}$, $\varepsilon_{3t} = u_{3t}$.

The properties of impulse response functions for nonlinear models are history dependent and are influenced by the sign and magnitude of the shocks. The data points of all variables must be chosen to illustrate the dynamic responses to shocks in different regimes. In our TAR models, the results show that the threshold variable of the excess stock returns equation is XR_{t-2}, the threshold variable of the change in the Federal funds rate equation is its own lag DFF_{t-2}, and the threshold variable of the growth rate of output equation is also XR_{t-2}, but with a different threshold value. There are six regimes if every possible situation is considered. In order to simplify the analysis and focus on the nonlinear effect of monetary policy on excess stock returns that we are interested, the data points of this paper are divided into two regimes by the threshold variable of the excess stock returns equation, XR_{t-2} and its threshold value -8.4. The data points in the low excess stock returns regime satisfy $XR_{t-2} \leq -8.4$ and the data points in the high excess stock returns regime satisfy $XR_{t-2} > -8.4$. The impulse response functions are calculated by averaging the impulse responses to shocks of 500 data points randomly drawn in each regime.

Figure 3.7 shows the impulse response functions of XR, DFF, and Gy for shocks to XR in the STAR model for both low excess returns and high excess returns regime. The impulse responses of XR to its own shocks are positive and fairly symmetric. The impulse responses of DFF to XR shocks show minor asymmetries in the sign of the shocks for both regimes. The responses of DFF to positive XR shocks have their own convergence patterns, and the responses of DFF to negative XR shocks

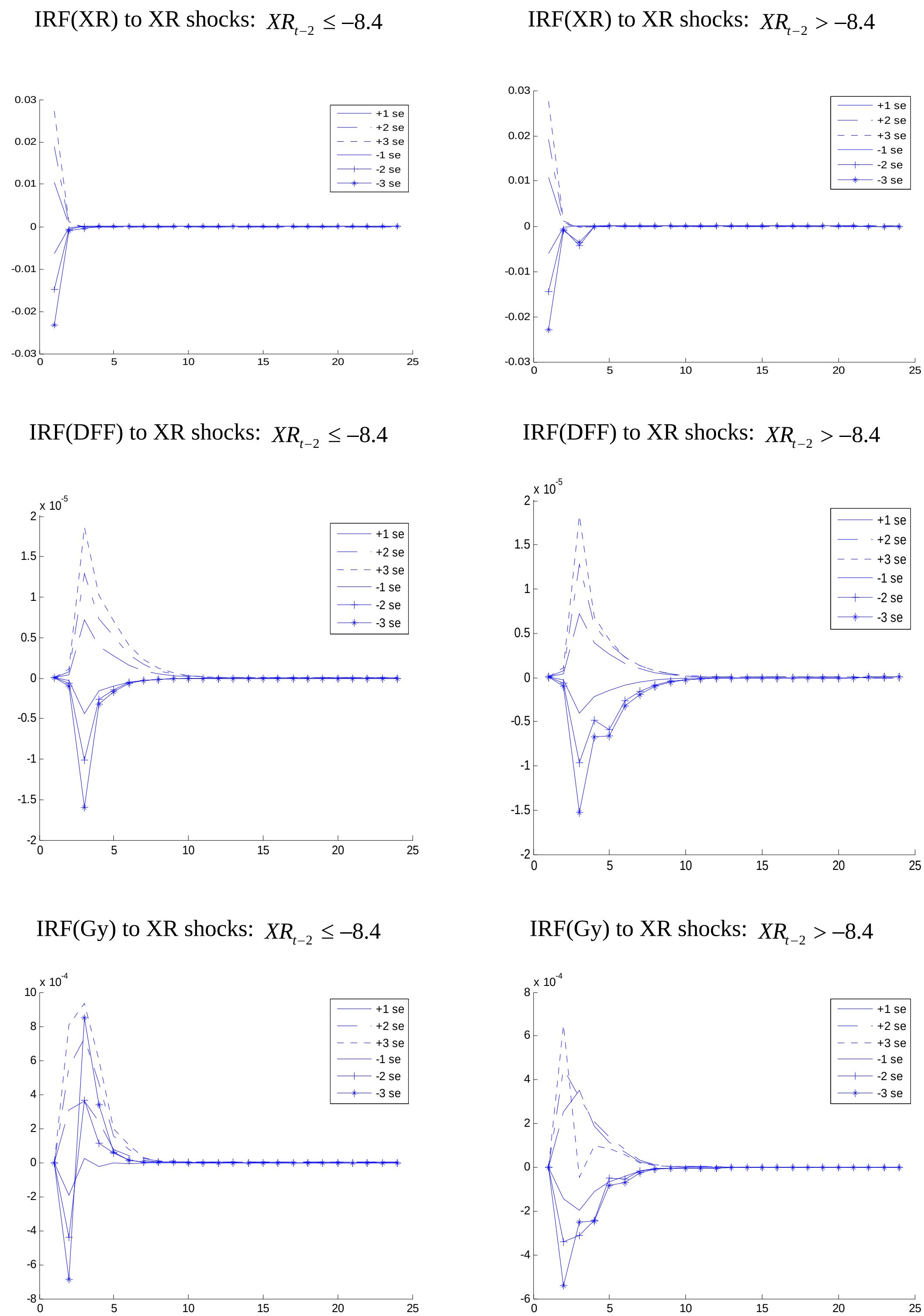

Figure 3.7. Impulse response functions for shocks to excess returns in two excess returns regimes

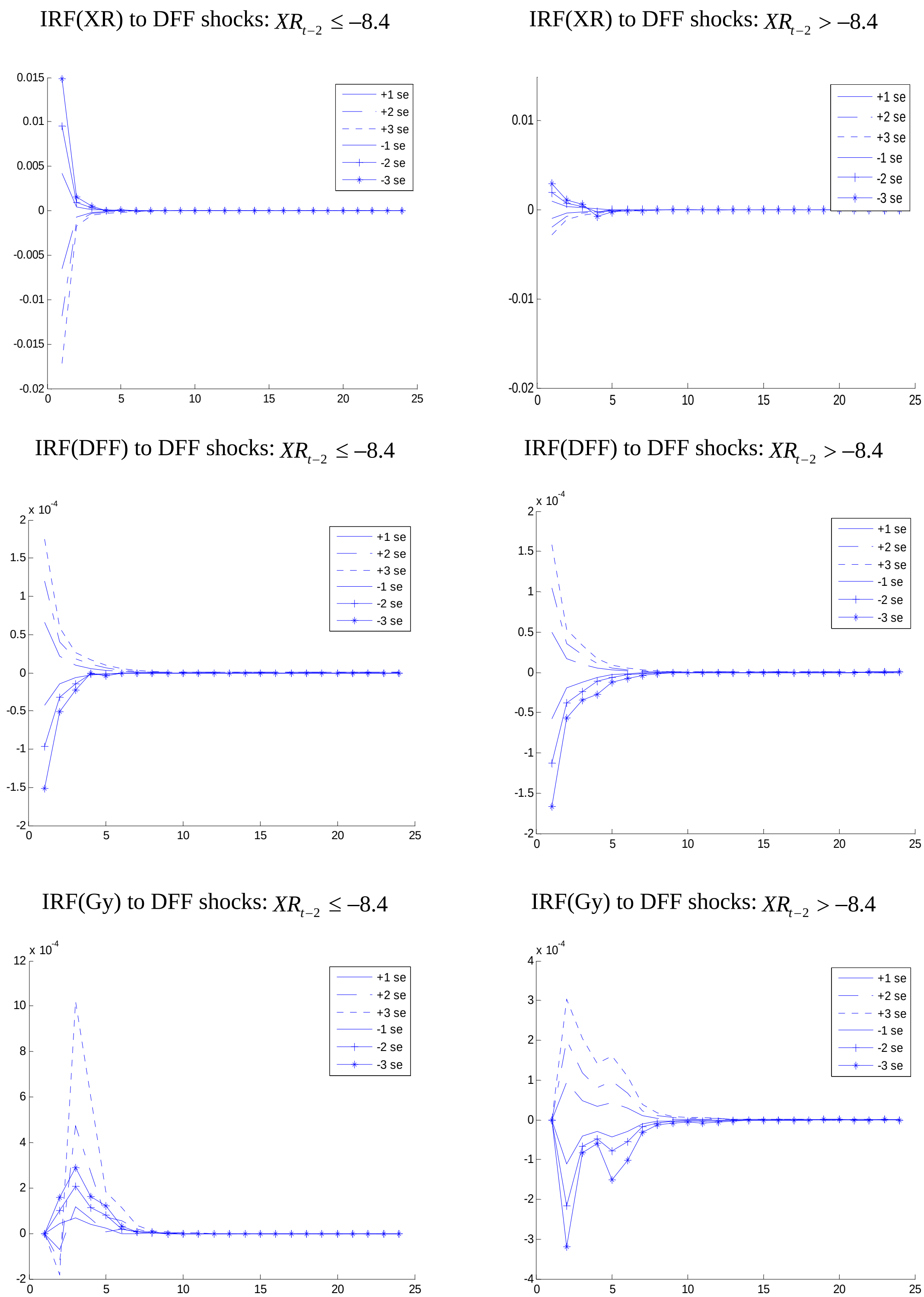

Figure 3.8. Impulse response functions for shocks to DFF in two excess returns regimes

have their own convergence patterns as well. Shocks to XR have positive effects on DFF and the size of the impulse responses function in both excess returns regimes are similar because the excess returns are not the transition variable that changes the parameters of DFF equation. The impulse responses of Gy to shocks to XR are really asymmetric and unstable after one period, especially the large negative shocks to XR in the low excess returns regime. It is because the excess return is also the transition variable of Gy equation. When there is a large negative shock to XR, it decreases Gy in the second period and then also decrease XR through the positive effects of Gy on XR in the excess returns equation. Once XR reduce to a extreme negative value, the negative effects of XR on Gy in the extreme excess returns regime will increase Gy.

Figure 3.8 shows the impulse response functions of XR, DFF, and Gy for shocks to DFF in the STAR model for $XR_{t-2} \leq -8.4$ and $XR_{t-2} > -8.4$ regime. The impulse responses of XR to DFF shocks are negative and the size is large in the low excess returns regime and very small in the high excess returns regime. This is consistent with the nonlinear coefficients in the STAR model of the excess return equation. The impulse responses of DFF to its own shocks are symmetric. The impulse responses of Gy to shocks to DFF in the high excess returns regime are positive, which is not consistent with our expectation. But, as we know from the results of model estimation, the effects of lags of DFF on Gy are not significant in the usual excess returns regime. The impulse responses of Gy to shocks to DFF in the low excess returns regime are unstable, again that is because the excess return is also the transition variable of Gy

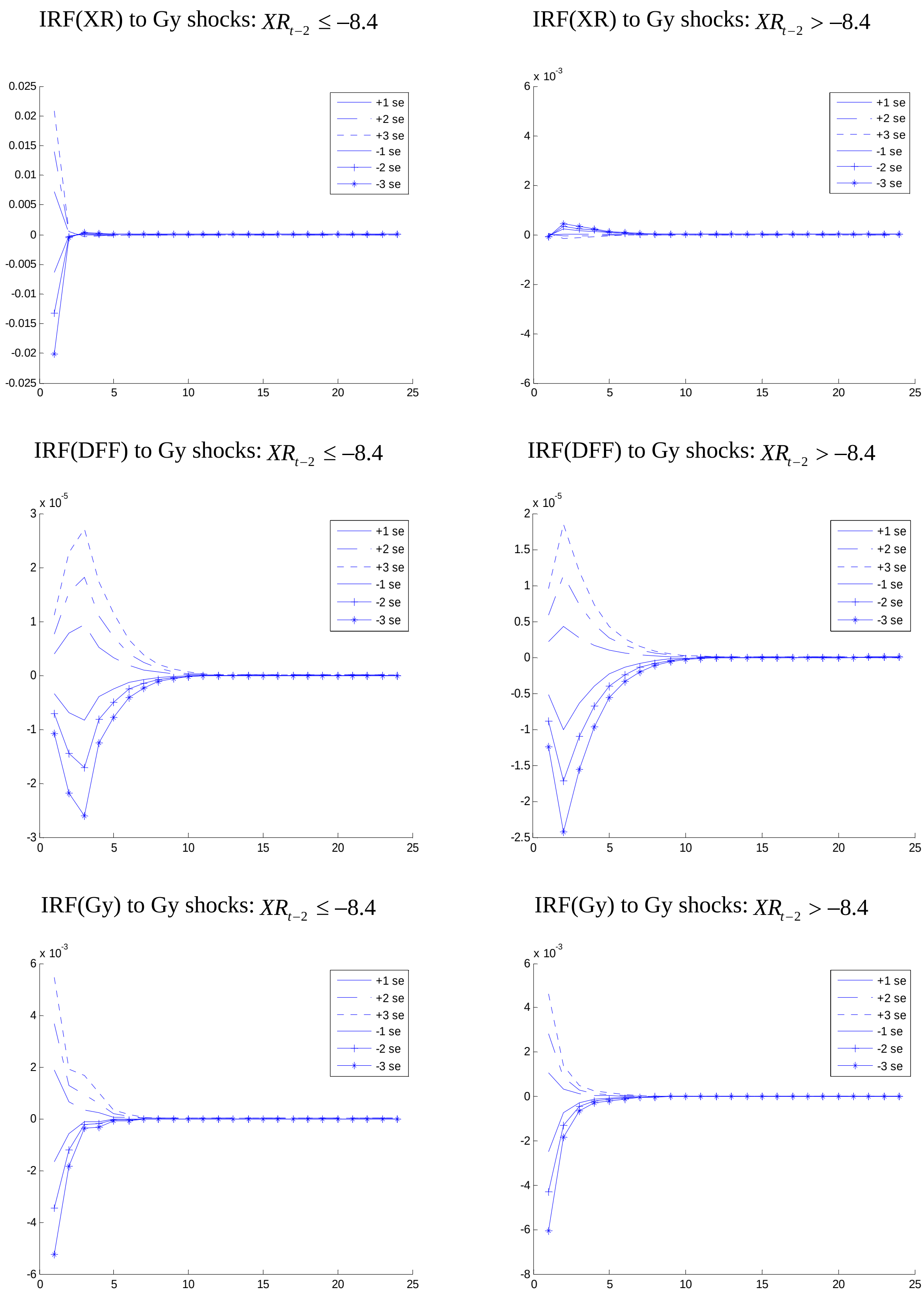

Figure 3.9. Impulse response functions for shocks to Gy in two excess returns regimes

equation. If a positive shock to DFF will decrease XR to the extreme excess returns regime of Gy equation, it will indirectly increase Gy to a positive number.

Figure 3.9 shows the impulse response functions of XR, DFF, and Gy for shocks to Gy in the STAR model for $XR_{t-2} \le -8.4$ and $XR_{t-2} > -8.4$ regime. The impulse responses of XR to shocks to Gy in the low excess returns regime are positive, but they are almost zero in the high excess returns regime. This nonlinear response shows that the growth rate of industrial production has larger impact on excess returns when excess return is pretty low. The impulse responses of DFF to shocks to Gy are positive and symmetric. The impulse responses of Gy to its own shocks are also positive and fairly symmetric.

3.6. Conclusion

In this paper I examine the impacts of monetary policy on excess returns of stock prices by using the smooth transition autoregressive (STAR) models. The change in the Federal funds rate is used as an endogenous measure of monetary policy and the growth rate of industrial production is also considered in the model. My empirical results show that excess stock returns, the change in the Federal funds rate, and the growth rate of industrial production all can be expressed in the nonlinear STAR models. The nonlinear dynamic model of excess stock returns is governed by the value of excess returns in two periods ago. The estimated coefficients and the impulse response functions show that the effect of monetary policy on excess returns of stock prices is significantly negative and nonlinear. The change in the Federal funds rate has a larger negative effect in the extreme low excess returns regime and has smaller effect when the excess return is greater than the threshold value.

CHAPTER IV

THE THRESHOLD EFFECTS OF EXCHANGE RATE VOLATILITY ON EXPORTS: EVIDENCE FROM U.S. BILATERAL EXPORTS

4.1 Introduction

The effect of exchange-rate volatility on trade flows has been much debated in the literature, especially after the collapse of the Bretton Woods system of fixed exchange rates in March, 1973. A number of studies have investigated this topic by adopting both theoretical models and empirical methods. From the theoretical point of view, the volatility of exchange rate can be seen as a risk in international trade. Hooper and Kohlhagen (1978) propose that under the assumption of risk aversion, an increase in exchange rate volatility increases the uncertainty of transactions, which leads to impede the volume of trade. On the contrary, if traders' preferences are risk love, exchange-rate volatility might increase the volume of trade.

De Grauwe (1988) argues that the impact of exchange rate volatility on exports depends on the degree of risk aversion, so even in the framework of risk aversion, a positive effect of exchange-rate volatility on exports can be seen. The economic intuition is that the very risk averse individuals worry a lot about the worse possible outcome, so when risk increases they will export more to avoid the possibility of a drastic decline in their revenues. Less risk averse individuals are less concerned with extreme outcomes. They view the return on export activity now less attractive when exchange rate risk increases and decide to export less.

De Grauwe also explains that an increase in exchange rate risk has a substitution effect and an income effect. The substitution effect is that an increase in exchange rate risk lowers the attractiveness of risky activities and leads to reduce exports. There is, however, also an income effect which works in the opposite direction. When exchange rate risk increases, the expected utility of export revenue declines, and this can be offset by increasing exports. Therefore, if the income effect is bigger than the substitution effect, the effect of exchange-rate volatility on exports is positive.

Arize et al. (2000) argues that because theory alone can not determine the sign of the relation between foreign trade and exchange rate volatility, the impact of exchange rate volatility on international trade is an empirical issue. They use time-series methodology which includes the cointegrating relations and the vector error-correction model (VECM) to estimate the long-run and short-run relationships between foreign trade and exchange rate volatility of thirteen less developed countries (LDC’s). Their results show that the increases in the volatility of the exchange rate have a significant negative effect on exports demand of LDC’s. Their conclusion is very close to the results of Chowdhury (1993) who investigates the G-7 countries by using VECM and get a significant negative impact of exchange-rate volatility on the volume of exports. Arize (1995, 1997) also employs export data of U.S. and seven industrial countries to test this issue empirically. Similarly, a negative relation between volume of export and exchange-rate volatility is the major result in those papers.

Comparing to lots of empirical research in this topic have been done by using aggregate foreign exports or imports data, bilateral trade data approach has been paid relatively less attention and the conclusions are still mixed. Koray and Lastrapes (1989) find that the effect of exchange-rate volatility is negative and weak on U.S. bilateral imports from U.K. Germany, France, Japan and Canada. Daly (1998) investigates the impact of exchange rate uncertainty on the export and import flow from Japan to her seven partner countries. His results show that the exchange rate volatility could raise the trade flows as well as reducing trade flows. Klaassen (2004) uses data on bilateral aggregate U.S. exports to the other G7 countries and argues that the effect is insignificant.

A few papers propose the possibility of nonlinear effects of exchange rate volatility on trade since there is no consistent conclusions in the literature. Baum et al. (2004) include foreign income volatility, by itself and in conjunction with the exchange rate volatility in the regression to capture the nonlinearities in the relationship between exchange rate uncertainty and bilateral exports of 13developed countries between 1980-1998. They find that the impact of exchange rate volatility on exports can intensify or diminish through changes in foreign income volatility. Zhang et al. (2006) test the threshold effect of exchange rate volatility on exports by employing time-series econometric techniques and bilateral export volumes data to the U.S. from the other six G7 countries. They find the existence of nonlinearity in the responses of export volumes to exchange rate volatility, and indicate that export volume tends to increase when exchange rate volatility surpasses a certain threshold point.

This paper wants to reexamine the effects of exchange rate volatility on export volume by using a new empirical method. There are three features of this paper that are different form previous empirical studies. First, instead of employing time-series technology that most papers use, a balanced panel data approach is adopted in this paper to test the effect of exchange rate volatility on the volume of bilateral export flows from U.S. to her thirteen major trade partners. The major advantage of using panel data is that it controls for time-invariant country heterogeneity. Dell'Ariccia (1999) is the only one who uses panel data to discuss the effects of exchange rate volatility on trade flows for Western Europe countries and exchange rate volatility is found to have a negative effect on international trade.

Second, we consider the possibility of nonlinear effects of exchange rate volatility on exports by using threshold regression methods for non-dynamic panels with individual-specific fixed effects proposed by Hansen (1999). Two variables are checked separately as the possible threshold variables. One is the bilateral real exchange rate volatility between U.S. and her importing partners. This is to confirm whether exchange rate volatility has the threshold effect on exports in Zhang et al. (2006). The other is the relative real gross domestic product (GDP) per capita of importing countries to U.S. According to the theoretical paper De Grauwe (1988), the impact of exchange rate volatility on exports depends on the degree of risk aversion and the very risk averse individuals will export more to avoid the possibility of a drastic decline in their revenues when risk increases. Also, from Arrow (1965) and McKee (1989), the hypothesis of decreasing absolute risk aversion says that as income rises, individuals are less risk averse to bets of fixed absolute size. These infer that lower income countries are more risk averse than higher income countries and might want to export more to avoid the possibility of a drastic decline in their revenues. Thus, income might be the threshold variable for explaining the nonlinear effects of exchange rate volatility on exports.

Finally, the thirteen major trade partners of U.S. in my sample cover developed countries and developing countries. It allows us to know if the country-specific character between developed and developing countries will be the reason to explain the nonlinear effects of exchange rate volatility on the volume of bilateral exports.

My empirical results show that exchange rate volatility has significant negative effect on the bilateral export volumes from U.S. to her thirteen major trading partners by using the linear panel data regression with fixed effects model. After testing the hypothesis of the threshold effect, evidence shows that exchange rate volatility has a nonlinear effect on U.S. bilateral exports only when the relative real GDP per capita to U.S. is the threshold variable. Exchange rate volatility reduces bilateral exports from U.S. to the country whose relative real GDP per capita to U.S. is greater than the threshold level, but increase bilateral exports from U.S. to the countries whose relative real GDP per capita to U.S are lower than the threshold level. This conclusion is robust when we estimate our model for top 30 major exporting partners of the United States. This confirms the inference that lower income countries are more risk averse and will increase exports when exchange rate volatility rises (income effects are larger than substitution effects); higher income countries are less risk averse and will reduce exports when exchange rate volatility rises (income effects are smaller than substitution effects).

The rest of the paper is organized as follows. Section 2 contains the model specifications of a linear panel data regression and a balanced panel threshold regression.

Section 3 describes the data and two measurements of exchange rate volatility. Section 4 reports the empirical results. Section 5 is the robustness check for top 30 major exporting partners of U.S. A brief conclusion is included in Section 6.

4.2 Model Specification

First, I estimate a static, non-dynamic linear panel data regression with individual-specific fixed effect model:

$$EX_{it} = \mu_i + \theta_1 Y_{it-1} + \theta_2 P_{t-1} + \theta_3 DEU_{it} + \beta_1 V_{it-1} + \beta_2 V_{it-1} Dfix_{it-1} + e_{it} \tag{4.1}$$

where the subscript i indexes the importing partner i of U.S. and the subscript t indexes time. EX_{it} denotes real bilateral export volume of U.S. to country i. μ_i denotes the unobservable individual specific effect. Y_{it-1} represents the real foreign economic activity for country i at time $t-1$, which is measured by the real GDP of country i. P_{t-1} represents the competitiveness of exporters and is measured by the ratio of U.S. export price to the world export price in U.S. dollar. DEU_{it} is the dummy variable of using Euro. $DEU_{it} = 1$, if $t \geq$ starting date of using Euro in country i; $DEU_{it} = 0$, if otherwise. V_{it-1} is the bilateral real exchange rate volatility between country i and U.S. Since few countries adopt the fixed exchange rate policy for some period of time in my sample, the real bilateral exchange rate volatility in the fixed exchange rate regime is actually from the volatility of relative price between U.S. and country i. In order to clearly understand that the effects of real exchange rate volatility on exports are from nominal exchange rate volatility, the dummy variable for fixed exchange rate regime

$Dfix_{it-1}$ is included in the regression. $Dfix_{it-1} = 1$, if country i adopts the fixed exchange rate policy at time $t-1$; $Dfix_{it-1} = 0$, if otherwise. Therefore, the effect of real exchange rate volatility on real bilateral exports of U.S. to country i is β_1 for float exchange rate periods and $\beta_1 + \beta_2$ for fixed exchange rate periods. e_{it} denotes the disturbance and is assumed to be independent and identically distributed with mean 0 and finite variance σ^2.

The coefficient θ_1 is expected to be positive because under the assumption of exporting normal goods, an increase in the income of importing country should raise the volume of exports. The coefficient θ_2 is expected to be negative since the demand of imports from U.S. should be reduced when relative export price of U.S. increases. The sign of the coefficient of DEU_{it} is ambiguous. The effect of real exchange rate volatility, β_1 for float exchange rate periods and $\beta_1 + \beta_2$ for fixed exchange rate periods, could be positive, negative or indeterminate, depending on different theories.

Next, I consider the possibility of nonlinear effects of real exchange rate volatility on bilateral exports. The non-dynamic balanced panel threshold regression with individual-specific fixed effect model is:

$$\begin{aligned} EX_{it} = {} & \mu_i + \theta_1 Y_{it-1} + \theta_2 P_{t-1} + \theta_3 DEU_{it} + \beta_{11} V_{it-1} I(q_{it-1} \le \gamma) \\ & + \beta_{12} V_{it-1} Dfix_{it-1} I(q_{it-1} \le \gamma) + \beta_{21} V_{it-1} I(q_{it-1} > \gamma) \\ & + \beta_{22} V_{it-1} Dfix_{it-1} I(q_{it-1} > \gamma) + e_{it} \end{aligned} \tag{4.2}$$

where $I(.)$ is the indicator function, q_{it-1} is the threshold variable, γ is the threshold value. There are two regimes in this model. When q_{it-1} is smaller than γ, the effect of real exchange rate volatility on real bilateral exports of U.S. to country i is β_{11} for float exchange rate periods and $\beta_{11}+\beta_{12}$ for fixed exchange rate periods, but when q_{it-1} is greater than γ, the effect of real exchange rate volatility on real bilateral exports of U.S. to country i is β_{21} for float exchange rate periods and $\beta_{21}+\beta_{22}$ for fixed exchange rate periods. Two possible threshold variables q_{it-1} considered in this paper are bilateral real exchange rate volatility V_{it-1} and relative real GDP per capita of importing countries i to U.S. Ypc_{it-1}.

Following Hansen (1999), for any given γ, the slope coefficients can be estimated by ordinary least squares estimation after fixed-effects transformations. The optimal threshold value $\hat{\gamma}$ is selected by the following procedure. Sort the distinct values of the observations on the threshold variable q_{it-1} and eliminate the smallest and largest 5% of the observations of threshold variable, so that at least 5% of the observations are in each regime. For the remaining γ, regressions are estimated and the optimal threshold value $\hat{\gamma}$ is the one which has the smallest sum of squared errors. For the hypothesis testing of the threshold effect, since under null hypothesis the threshold γ is not identified, classical tests have non-standard distributions. Hansen (1999) suggests a bootstrap to simulate the asymptotic distribution of the likelihood ratio test. More details of the procedures and mathematical proofs are described in Hansen (1999).

4.3 Data

All data are taken from International Financial Statistics (IFS) data bank and Directions of Trade (DOT). The countries included are thirteen major trading partners of U.S.: Canada, Mexico, Japan, U.K., Belgium, Germany, Korea, Netherlands, France, Singapore, Australia, Italy, and Malaysia, which cover about 75% of U.S. exports.[2] The order of magnitude of the bilateral export shares from first major partner to the thirteenth is the same as the order in which they are listed above. The frequency of data is quarterly. The sample period is 1973:2~2004:4. Our sample period starts from 1973:2 because the abandonment of fixed exchange rates for most of the countries was in March 1973. The real bilateral exports (EX_{it}) is the natural logarithm of bilateral export volume of U.S. to country i (2000=100). The real economic activity of importing country i (Y_{it-1}) is measured by the natural logarithm of GDP volume of country i (2000=100).[3] The relative price of exports (P_{t-1}) is defined as the natural logarithm of the ratio of U.S. export price to the world export price in U.S. dollar. The relative real GDP per capita of country i to U.S. (Ypc_{it-1}) is the natural logarithm of the ratio of real GDP per capita of country i expressed in U.S. dollar to real GDP per capita of U.S. (The real GDP per capita is deflated by local consumer price index).

[2] Since the data on China, Brazil, and Hong Kong is not available in our sample period, we exclude those countries in our study.
[3] The quarterly data of GDP volume might not available from 1973:2 for some countries, so the annual data of GDP volume are converted to a quarterly basis by using a quadratic interpolation method proposed by Goldstein and Khan (1978) for the lack data period.

There are two popular methods to measure the real exchange rate volatility V_{it} in the literature. One is the moving sample standard deviation of the real exchange rate changes

$$V_{i,t+m}^{MS} = [\frac{1}{m}\sum_{k=1}^{m}(R_{i,t+k} - R_{i,t+k-1})^2]^{1/2} \tag{4.3}$$

where m = 8 is the order of moving average; R_{it} denotes the natural logarithm of real bilateral exchange rate (rer) between U.S. and country i, and rer is

$$reer(i,t) = \exp[\ln E(us,\$,t) - \ln p(us,t) - \ln E(i,\$,t) + \ln p(i,t)] \tag{4.4}$$

where $E(i,\$,t)$ is the exchange rate in units of i currency per U.S. dollars in index form (2000=100)[4]; p denotes the consumer price index of country i (2000=100). The other method of measuring the real exchange rate volatility is to obtain the conditional standard deviation, which is denoted as V_{it}^{GARCH}, from a GARCH (1,1) based on an autoregressive model of order 1 of the first difference of R_t for each country.

$$DR_t = \alpha_0 + \alpha_1 DR_{t-1} + u_t,\ \ u_t \mid I_{t-1} \sim N(0, v_t^2) \tag{4.5}$$

$$v^2 = h_0 + h_1 u_{t-1}^2 + h_2 v_{t-1}^2 \tag{4.6}$$

where DR_t is the first difference of the logarithm of the real bilateral exchange rate, v_t^2 is the conditional variance of the error term u_t. Eight observations at the beginning of the sample are lost due to the calculation of the volatility measure V_{it}^{MS} and two

[4] Because of the circulation of Euro, it has a sudden drop or jump in the data of exchange rate from 1999. In order to keep the same exchange rate unit, we transfer the new exchange rate unit (Euro per U.S. dollar) back to the original exchange rate unit (currency of country i per U.S. dollar) for the member of European Union by using the official rate between Euro and currency of country i.

observations at the beginning of the sample are lost due to the calculation of V_{it}^{GARCH}, so the estimation periods for using V_{it}^{MS} and V_{it}^{GARCH} as the exchange rate volatility are 1975:2~2004:4 and 1973:4~2004:4, respectively.

4.4 Empirical Results

First we estimate the linear panel data regression model for the bilateral exports from U.S. to her thirteen major trading partners. The results of linear panel data regression model are presented in Table 4.1. Column 2 shows the estimated coefficients for measuring the exchange rate volatility by using moving sample standard deviation method ($V_{it} = V_{it}^{MS}$) and column 3 shows the result of measuring exchange rate volatility from a GARCH(1,1) model ($V_{it} = V_{it}^{GARCH}$). The effects of real economic activity of importing countries are all positive and statistically significant for two kinds of exchange rate volatility models. The responses of bilateral exports to relative price of exports are all negative and statistically significant. Those results are all consistent with the expectation. The coefficient of dummy variable DEU_{it} is significantly positive, which means that adopting Euro grows the volume of exports from U.S. to her trading partners in EU. The estimated coefficients of exchange rate volatility V_{it-1} that we are interested are negative and statistically significant for both cases. This implies that an increase in the real exchange rate volatility reduces the bilateral exports. The estimated coefficients of $V_{it-1} * Dfix_{t-1}$ are negative and significant, which means the negative effects of real exchange rate volatility on bilateral exports become larger in the fixed exchange rate periods.

Table 4.1 Linear panel data regression estimates

	if $V_{it-1} = V_{it-1}^{MS}$	if $V_{it-1} = V_{it-1}^{GARCH}$
Y_{it-1}	1.413 (0.016)	1.390 (0.016)
P_{t-1}	-0.701 (0.087)	-0.717 (0.088)
DEU_{it}	0.074 (0.023)	0.088 (0.024)
V_{it-1}	-2.238 (0.279)	-1.150 (0.270)
$V_{it-1} * Dfix_{t-1}$	-2.216 (0.556)	-2.499 (0.470)
R^2	0.849	0.853

Note: Standard errors are in parentheses.

Before estimating the panel threshold regression model, we need to determine whether a threshold relationship exists between real bilateral exchange rate volatility and real bilateral exports. According to Hansen (1999), we adopt the likelihood ratio test and construct p-values of the test statistics by bootstrap. The Null hypothesis is no threshold effect and the alternative hypothesis is one threshold effect. The testing results are shown in Table 4.2. When exchange rate volatility V_{it-1} is the threshold variable, the test statistics F_1 can not be rejected under the null hypothesis. Exchange-rate volatilities V_{it-1} have no threshold effect for the impacts of real exchange rate volatility on bilateral exports. But when the threshold variable is the relative real GDP per capita of country i to U.S. (Ypc_{it-1}), F_1 are highly significant with a bootstrap p-value less than 5%. So,

Ypc_{it-1} has threshold effect for the impacts of real exchange rate volatility on bilateral exports.

Table 4.2 Test for threshold effects

	if $V_{it-1} = V_{it-1}^{MS}$		if $V_{it-1} = V_{it-1}^{GARCH}$	
q_{it-1}	V_{it-1}	Ypc_{it-1}	V_{it-1}	Ypc_{it-1}
Test for a single threshold				
F_1	31.006	480.411	81.651	518.964
P-value	0.754	0.004	0.354	0.002
(10%, 5%, 1% critical values)	(76.263)	(142.091)	(194.201)	(170.435)
	(91.431)	(249.497)	(266.222)	(256.617)
	(133.470)	(422.314)	(409.704)	(360.196)

Note: Standard errors are in parentheses.

The estimates of panel threshold regression model with the threshold variable Ypc_{it-1}, the relative real GDP per capita to U.S., are displayed in Table 4.3. The second column shows the estimated coefficients when exchange rate volatility is measured by moving sample standard deviation (V_{it} = V_{it}^{MS}). Again, the real GDP Y_{it} has a positive and significant effect on bilateral exports. The relative price P_t which represents the competitiveness has a significant negative effect. Those results are consistent with the theoretical thinking and the effect of each variable is similar to the linear panel data regression model in Table 4.1. The coefficient of dummy variable DEU_{it} is positive and significant as before.

The estimated threshold value γ is -1.270. We then define the economy of importing country i is in low income regime if Ypc_{it} is less than -1.262 and is in high income regime if Ypc_{it} is greater than -1.270. For the float exchange rate period, the coefficient of V_{it-1} is positive and statistically significant in low income regime, but becomes negative and statistically significant in high income regime. This implies that when the income of importing country i is far below the income of U.S. ($Ypc_{it-1} \leq -1.270$), an increase in the exchange rate volatility induces the importing country to increase its import activity, but when the relative income of importing country i to U.S. is greater than the threshold ($Ypc_{it-1} > -1.270$), a higher exchange rate volatility leads the importing country to import less.

Table 4.3 Estimates of panel threshold regression model

	if $V_{it-1} = V_{it-1}^{MS}$	if $V_{it-1} = V_{it-1}^{GARCH}$
Y_{it-1}	1.441 (0.014)	1.440 (0.016)
P_{t-1}	-0.700 (0.076)	-0.755 (0.077)
DEU_{it}	0.050 (0.020)	0.067 (0.017)
$V_{it-1}I(Ypc_{it-1} \leq \gamma)$	1.541 (0.320)	1.574 (0.770)
$V_{it-1}Dfix_{it-1}I(Ypc_{it-1} \leq \gamma)$	2.185 (0.813)	2.935 (1.270)
$V_{it-1}I(Ypc_{it-1} > \gamma)$	-5.295 (0.292)	-5.544 (0.518)
$V_{it-1}Dfix_{it-1}I(Ypc_{it-1} > \gamma)$	-3.028 (0.622)	-0.567 (1.462)
γ	-1.270	-1.288
R^2	0.886	0.889

Note: Standard errors are in parentheses.

These results confirm the inference that comes from De Grauwe (1988) and Arrow (1965). De Grauwe (1988) proposes that if the traders are sufficiently risk averse, an increase in the exchange rate volatility raises the expected marginal utility of revenue and therefore induces them to increase their trading activity, i.e. the positive income effect dominates the negative substitution effect, traders will export more to avoid the possibility of a drastic decline in their revenues when exchange rate risk increases. The impact of exchange rate volatility on exports depends on the degree of risk aversion. Arrow (1965) generates the hypothesis of decreasing absolute risk aversion. As income rises, individuals are less risk averse to bets of fixed absolute size. These explain that lower income countries are more risk averse than higher income countries and want to import more to avoid the possibility of a drastic decline in their revenues when real exchange rate volatility occurs. Thus, income has the threshold effect for the impacts of real exchange rate volatility on bilateral exports. For the fixed exchange rate periods, the source of the real exchange rate volatility is from the volatility of relative price between U.S. and country *i* and the results show that the impacts of the real exchange rate volatility on exports become larger in absolute value for both income regimes.

The third column of estimates in Table 4.3 presents the results for the panel threshold model with real exchange rate volatility V_{it}^{GARCH} from a GARCH(1,1) model. The estimated threshold value γ is -1.288. All the coefficients are similar to those in the second column, except the real exchange rate volatility effects are not different between float and fixed exchange rate policy periods in the higher income regime. Table 4.4 reports the number of observations per income threshold regime and for different exchange rate policy periods.

In order to see what income regime the importing countries in my sample fall over time, Table 4.5 reports the percentage of observations in higher income regime by country. All developed countries (Australia, Belgium, Canada, France, Germany, Italy, Japan, Netherlands, and U.K.) are 100% falling in the high income regime. According to our nonlinear regression results, when the bilateral real exchange rate volatility V_{it} increases in those countries, the bilateral exports of U.S. to them decrease. Malaysia falls in the low income regime for all data period. An increase in V_{it} induces the bilateral exports of U.S. to Malaysia. Other developing countries (Korea, Mexico, and Singapore) fall in the high income regime only in part of data period, so the effects of real exchange rate volatility on exports of U.S. to them depend on which income regime they are in. If a country is in the low income regime, the exports of U.S. to the country increase as bilateral real exchange rate volatility increases. But when the relative real GDP per capita of the importing country to U.S. surpasses the threshold value, an increase in V_{it} will reduce the bilateral exports of U.S. to it.

Table 4.4 Number of observations by threshold and the exchange rate policy

Number of observations		if $V_{it-1} = V_{it-1}^{MS}$	if $V_{it-1} = V_{it-1}^{GARCH}$
$Ypc_{it-1} \leq \gamma$	Float exchange rate regime	195	229
	Fixed exchange rate regime	43	30
$Ypc_{it-1} > \gamma$	Float exchange rate regime	1285	1322
	Fixed exchange rate regime	24	44

Table 4.5 Percentage of observations in the high relative income regime by country

V_{it-1}	V_{it-1}^{MS}	V_{it-1}^{GARCH}
γ	-1.270	-1.251
Australia	100	100
Belgium	100	100
Canada	100	100
France	100	100
Germany	100	100
Italy	100	100
Japan	100	100
Korea	51.26	46.40
Malaysia	0	0
Mexico	66.39	68.00
Netherlands	100	100
Singapore	82.35	78.40
U.K.	100	100

4.5 Robustness Test

Since most of the major exporting partners of U.S. are developed countries, only four developing countries in our sample fall in the low income state for some sample periods. In order to check the robustness of the nonlinear effect of real exchange rate volatility on bilateral exports of U.S., I re-estimate the model by adding more trading partners of U.S. to the sample. I consider top 30 major exporting partners of U.S. which covers about 85% of total export volume of U.S., but due to lack of data for some countries only 23 countries are included.[5] Except the thirteen countries that we have used in the previous estimations, the rest ten countries are Saudi Arabia, Switzerland, Israel, Spain, Philippines, Thailand, Ireland, India, Colombia, and Sweden, which are

[5] The excluded countries are China, Brazil, Hong Kong, Venezuela, Rep. Bol., Russia, Argentina, and Egypt.

listed in the order of magnitude of the bilateral export shares.[6] We then test whether the threshold effect is statistically significant and estimate the appropriate panel regression models.

Table 4.6 presents the hypothesis testing results with threshold variable q_{it-1} = Ypc_{it-1} and q_{it-1} = V_{it-1} for 23 countries. No matter what kinds of exchange rate volatility V_{it-1} are used, when real exchange rate volatility V_{it-1} is the threshold variable, the test statistics F_1 can not be rejected under the null hypothesis, but the test F_1 is highly significant under 95% significant level when Ypc_{it-1} is the threshold variable. Thus, Ypc_{it-1} has threshold effect for the impacts of real exchange rate volatility on bilateral exports.

Table 4.6 Test for threshold effects for 23 countries

	if $V_{it-1}=V_{it-1}^{MS}$		if $V_{it-1}=V_{it-1}^{GARCH}$	
q_{it-1}	V_{it-1}	Ypc_{it-1}	V_{it-1}	Ypc_{it-1}
Test for a single threshold				
F_1	56.129	255.106	75.683	226.579
P-value	0.328	0.024	0.438	0.028
(10%, 5%, 1% critical values)	(92.624)	(181.044)	(159.262)	(184.796)
	(113.725)	(212.731)	(199.174)	(215.194)
	(187.022)	(314.871)	(281.480)	(284.996)

Note: Standard errors are in parentheses.

[6] A quadratic interpolation method is used to convert the annual data of GDP to a quarterly basis if the quarterly data of GDP are not available from 1973:2.

The estimates of one threshold panel regression models and the corresponded standard errors with threshold variable Ypc_{it-1} for 23 countries are shown in Table 4.7. The coefficients of Y_{it-1}, P_{t-1}, and DEU_{it} have the expected signs and similar magnitude as in Table 4.3. The threshold value γ is -1.528 and -1.595 when V_{it-1} = V_{it-1}^{MS} and V_{it-1} =

Table 4.7 Estimates of panel threshold regression model for 23 countries

	if $V_{it-1} = V_{it-1}^{MS}$	if $V_{it-1} = V_{it-1}^{GARCH}$
Y_{it-1}	1.354 (0.014)	1.344 (0.015)
P_{t-1}	-0.774 (0.071)	-0.847 (0.072)
DEU_{it}	0.067 (0.021)	0.092 (0.018)
$V_{it-1}I(Ypc_{it-1} \le \gamma)$	2.213 (0.328)	4.033 (0.630)
$V_{it-1}Dfix_{it-1}I(Ypc_{it-1} \le \gamma)$	0.154 (0.928)	0.300 (0.848)
$V_{it-1}I(Ypc_{it-1} > \gamma)$	-3.175 (0.282)	-1.778 (0.842)
$V_{it-1}Dfix_{it-1}I(Ypc_{it-1} > \gamma)$	-3.286 (0.655)	-2.451 (0.998)
γ	-1.528	-1.595
R^2	0.811	0.808

V_{it-1}^{GARCH}, respectively. For the float exchange rate periods, the estimates of the responses of bilateral exports to exchange rate volatility are positive and significant in the low income regime, but become significant negative in the high income regime. For the fixed exchange rate periods, the coefficients of $V_{it-1}Dfix_{it-1}I(Ypc_{it-1} \le \gamma)$ are not significant different from zero in the lower income regime, which implies that the impacts of real exchange rate volatility are not different for two kinds of exchange rate policies. But in the higher income regime, as the results of 13 major exporting partners, the responses of exports to real exchange rate volatility are bigger in the fixed exchange rate periods. Thus, the nonlinear effect of exchange rate volatility on bilateral exports with the threshold variable Ypc_{it-1} is confirmed again in this robustness check.

Table 4.8 reports the number of observations per income threshold regime and for different exchange rate policy periods and Table 4.9 shows the percentage of observations in higher income regime by country for 23 major exporting countries. All the developed countries are in the high income regime for the whole sample period. It indicates that an increase in the real exchange rate volatility will reduce the bilateral exports of U.S. to those developed countries. But, for other countries, Saudi Arabia and

Table 4.8 Number of observations by threshold and the exchange rate policy for 23 countries

Number of observations		if $V_{it-1} = V_{it-1}^{MS}$	if $V_{it-1} = V_{it-1}^{GARCH}$
$Ypc_{it-1} \le \gamma$	Float exchange rate regime	535	545
	Fixed exchange rate regime	65	64
$Ypc_{it-1} > \gamma$	Float exchange rate regime	2029	2136
	Fixed exchange rate regime	108	130

Singapore have 94.42% to 80% observations in the high income regime. Colombia, Korea, and Mexico have about 68.8% to 55.08% observations in the high income regime. Philippines has only 23.2% in the high income regime. India, Malaysia, and Thailand have all the observations in the low income regime. So, the effect of real exchange rate volatility on bilateral exports of U.S. to those developing countries depends on which regime the exporting partner will be in. The effect is positive if the exporting partner's real GDP per capita relative to that of U.S. is lower than the threshold value, but the effect is negative if the exporting partner's real GDP per capita relative to that of U.S. is greater than the threshold value.

Table 4.9. Percentage of observations in the high relative income regime by country for 23 countries

V_{it}	V_{it}^{MS}	V_{it}^{GARCH}
γ	-1.528	-1.595
Australia	100	100
Belgium	100	100
Canada	100	100
Colombia	55.46	58.40
France	100	100
Germany	100	100
India	0	0
Ireland	100	100
Israel	100	100
Italy	100	100
Japan	100	100
Korea	67.23	68.80
Malaysia	0	0
Mexico	68.91	72.00
Netherlands	100	100
Philippines	21.01	28.00
Saudi Arabia	94.96	100
Singapore	88.24	85.60
Spain	100	100
Sweden	100	100
Switzerland	100	100
Thailand	0	0
U.K.	100	100

4.6 Conclusion

In the previous literature, the effect of real exchange rate volatility on exports has been fully discussed by using time series data, but the conclusion is still mixed, especially using the bilateral exports data. In this paper, I attempt to reexamine the effects of real exchange rate volatility on U.S. bilateral export flows by using panel data approach which takes the time-invariant country heterogeneity into account. Moreover, I also consider the possibility of nonlinear effects of exchange rate volatility on exports by using threshold regression methods for non-dynamic panels with individual-specific fixed effects proposed by Hansen (1999).

The bilateral exchange rate volatility is measured by using moving sample standard deviation method and the conditional standard deviation from a GARCH (1,1) model. The data of real bilateral export volume from U.S. to her thirteen major trading partners are used. After estimating the regular linear panel data regression, the results show that exchange rate volatility has negative impact on exports, which is consistent with most of the empirical literature. However, the hypothesis of no threshold effect has been rejected when the relative real GDP per capita of the trading partner to U.S. is the threshold variable. The estimates from panel threshold regression model show that real exchange rate volatility has positive and statistically significant effect on bilateral exports in the low income regime, but when the relative real GDP per capita of the trading partner surpasses the threshold value, an increase of exchange rate volatility will reduce the bilateral exports of U.S. to it. This conclusion is robust when the model has been estimated again for top 30 major exporting partners of the United States.

CHAPTER V

CONCLUSIONS

In chapter II, the asymmetric effect of monetary policy on stock price is investigated in the framework of an unobserved-components model with Markov-switching heteroscedasticity (UC-MS model). I assume that monetary policy can only influence the transitory component of stock prices. By estimating the UC-MS model without monetary policy and the UC-MS model with monetary policy, my results show that monetary policy has negative effects on stock prices. A contractionary monetary policy significantly reduces stock prices in the low volatility state of the transitory component. When the transitory component is in the high volatility state, the negative effect of monetary policy becomes larger, but the difference of the monetary policy effects between two states is not significant. I also find that monetary policy can reduce the total volatility of stock prices and the volatility of the transitory component. Monetary policy can also affect the dynamics of switching between low volatility and high volatility state. A contractionary monetary policy will lower the probability of staying in the low volatility state.

Chapter III discusses the impacts of monetary policy on excess returns of stock prices by using the smooth transition autoregressive (STAR) models. My empirical results show that excess stock returns, the change in the Federal funds rate, and the growth rate of industrial production all can be expressed in the nonlinear STAR models. The nonlinear model of excess stock returns is controlled by its own lag at two. A contractionary monetary policy significantly reduces excess stock returns and the effect is nonlinear. The change in the Federal funds rate has a larger negative effect in the extreme low excess returns regime and has smaller effect when the excess return is greater than the threshold value.

Chapter IV tries to answer the mix results for the topic of the effect of real exchange rate volatility on exports. I examine the effect of real exchange rate volatility on U.S. bilateral export flows by using panel data approach which takes the time-invariant country heterogeneity into account. I also consider the possibility of nonlinear effects of exchange rate volatility on exports by using threshold regression methods for non-dynamic panels with individual-specific fixed effects proposed by Hansen (1999). The results indicate that exchange rate volatility has nonlinear impact on bilateral exports. Real exchange rate volatility has positive and statistically significant effect on bilateral exports in the low income regime, but when the relative real GDP per capita of the trading partner surpasses the threshold value, an increase of exchange rate volatility will reduce the bilateral exports of U.S. to it. This conclusion is robust when the model has been estimated again for top 30 major exporting partners of the United States.

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