ESSAYS ON MONETARY ECONOMICS AND FINANCIAL ECONOMICS

A Dissertation

by

SOK WON KIM

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

August 2006

Major Subject: Economics

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Approved by:

Chair of Committee, Dennis Jansen
Committee Members, Li Gan
Byeongseon Seo
Kishore Gawande
Head of Department, Amy Glass

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ABSTRACT

Essays on Monetary Economics and Financial Economics.

(August 2006)

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Chair of Advisory Committee: Dr. Dennis Jansen

In this dissertation three different economic issues have been analyzed. The first issue is whether monetary policy rules can improve forecasting accuracy of inflation. The second is whether the preference of a central bank is symmetry or not. The last issue is whether the behavior of aggregate dividends is asymmetry. Each issue is considered in Chapter II, III and IV, respectively.

The linkage between monetary policy rules and the prediction of inflation is explored in Chapter II. Our analysis finds that the prediction performance of the term structure model hinges on monetary policy rules, which involve the manipulation of the federal funds rate in response to the change in the price level. As the Fed's reaction to inflation becomes stronger, the predictive information contained in the term structure becomes weaker. Using the long-run Taylor rule, a new assessment of the prediction performance regarding future change in inflation is provided. The empirical results indicate that the long-run Taylor rule improves forecasting accuracy.

In chapter III, the asymmetric preferences of the central bank of Korea are examined under New Keynesian sticky prices forward-looking economy framework. To this end,

this chapter adopts the central bank's objective functional form as a linear-exponential function instead of the standard quadratic function. The monetary policy reaction function is derived and then asymmetric preference parameters are estimated during the inflation targeting period: 1998:9-2005:12. The empirical evidence supports that while the objective of output stability is symmetry, but the objective of price stability is not symmetry. Specifically, it appears that the central bank of Korea aggressively responds to positive inflation gaps compared to negative inflation gaps.

Chapter IV examines the nonlinear dividend behavior of the aggregate stock market. We propose a nonlinear dividend model that assumes managers minimize the regime dependent adjustment costs associated with being away from their target dividend payout. By using the threshold vector error correction model, we find significant evidence of a threshold effect in aggregate dividends of S&P 500 Index in quarterly data when real stock prices are used for the target. We also find that when dividends are relatively higher than target, the adjustment cost of dividends is much smaller than that when they are lower.

To Mehye and Eunho

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A very special note of thanks is due to my wife and daughter. Without their patience and love I could not have finished this long and demanding journey. Thanks also go to my families who live in Korea for their endless support.

TABLE OF CONTENTS

CHAPTER Page

LIST OF TABLES

LIST OF FIGURES

CHAPTER I

INTRODUCTION

The objective of this dissertation is to analyze three different economic issues: inflation forecasting with the monetary policy rule, an asymmetric preference of central bank, and an asymmetric adjustment in aggregate dividends. Each issue is covered in Chapter II, III and IV, respectively. The main findings are summarized in Chapter V.

With regard to inflation forecasting, we consider a term structure model of inflation forecasting in Chapter II. The rational expectations model implies that asset prices reflect forward-looking behavior in the financial market. In particular, the term structure of interest rates provides potential information on the prediction of interest rates and inflation according to the expectations hypothesis and the Fisher equation. The predictive information contained in the yield curve has been analyzed in many empirical studies. The empirical results show that the prediction performance of the term structure model varies depending on the maturities of the yield curve and the sample period. Mishkin (1990) has shown that the term structure provides almost no information about the future change in inflation for maturities of six months or less.

It is natural to ask what affects the prediction performance of the term structure model. It is noteworthy that the term structure of interest rates reveals the stylized facts of temporal persistence as discussed in Seo (2003) compared to the variation of the change

This dissertation follows the style of Journal of Monetary Economics.

in inflation. The stylized facts indicate imbalance in regression. Many studies have shown that the persistence of the term spread is related to the monetary policy. For example, Mankiw and Miron (1986) provided empirical results that the predictive information of the term structure began to disappear after the founding of the Federal Reserve and its manipulation of interest rates. Necessarily, the variation in inflation is associated with the Fed's reaction to inflation.

Another important issue in forecasting inflation is associated with parameter instability. The Phillips curve relates the unemployment rate to a measure of inflation. Thus, the Phillips-curve-based inflation forecasts have been used widely in monetary policymaking. However, these forecasts have been found to be sensitive to instability, particularly in the 1970s and early 1980s.

Although there is a vast literature on the monetary policy rules, there have been no attempts to relate the monetary policy rules to the prediction of inflation. In this chapter, we consider the inflation forecasts using the monetary policy rules. As the monetary policy rules may differ between the monetary policy regimes, we examine the parameter stability by using the statistical methods.

In this chapter, using the U.S. monthly data for the period January 1960-December 2004, an empirical assessment of the linkage between the monetary policy rules and the prediction of inflation is provided. As the rational expectations model does not consider the effect of the monetary policy rules, this study resolves the mismatch between economic theory and empirical findings.

In Chapter III, the preference of central banks is analyzed under their asymmetric objective function. In analyzing their optimal monetary policy or monetary policy rules, traditionally literature has assumed that the preferences of the central banks are symmetric over key macroeconomic variables, such as inflation and output. However, recently a growing number of papers have questioned this symmetric preference assumption: specifically, are negative deviations of inflation from target as undesirable as positive deviations of inflation from target of the same amount? And/or are positive output gaps as distasteful as negative output gaps of the same size?

The asymmetric loss function generally leads to a nonlinear reaction function or monetary policy rule which is the first order condition solved for the optimization problem of central banks. Whether the central bank has the asymmetric objective function is an important issue since many of the results on the time consistency problem under symmetric preferences may no longer hold under asymmetric preferences as shown in Nobay and Peel (2003), and Surcio (2003a).

Recently the way of conducting monetary policy in Korea was changed dramatically. After the Bank of Korea Act revised in April 1998, the inflation targeting was adopted as new monetary policy system. In addition the Bank of Korea, which is the central bank in Korea, uses interest rate as its monetary policy instrument instead of monetary aggregates. Since the Bank of Korea has set up the explicit inflation target from 1998, the study on its asymmetric preference of inflation will have an advantage over assuming the implicit inflation target frequently used for the non-inflation targeting countries, for example United States.

The goal of this chapter examines the possible asymmetric preference of the central bank in Korea during the inflation targeting period: 1998:9-2005:12. Specifically, we try to answer the question whether the preferences for inflation or/and output gap of the central bank in Korea are asymmetry or not.

Chapter IV investigates the asymmetric behavior of aggregate dividends. After Lintner (1956) proposed the well-known behavioral model of dividend policy, a number of papers try to model the behavior of dividend in both disaggregate and aggregate level. For aggregate level, Marsh and Merton (1987) developed the dynamic behavior model of aggregate dividend which has a feature that dividend is adjusted to its long run target dividend (i.e., error correction term). Garrett and Priestley (2000) generalize Lintner model by introducing the manager's optimizing behavior. Most literature has analyzed the dividend behavior in the linear functional form.

However, there is some possibility that adjustment of dividends may not be a linear process. An important feature of dividends behavior is that there is asymmetry in dividend payments due to, for instance, a reluctance to cut dividends. For example, Yoon and Starks (1995) document the evidence that there is an asymmetry between dividend increases and dividend decreases at the individual firm level. Jalilvand and Harris (1984) examined the process of partial adjustment by allowing speeds of adjustment to vary by firm and over time depending on the size of firm and capital market conditions such as interest rates and stock prices. In addition, Marsh and Merton (1987) supported asymmetric adjustment of dividends.

Therefore, we examine whether adjustment cost of dividends depends on the state of previous difference between dividends and the target. For example, adjustment cost of dividends faced by managers may be low (or high) when the previous dividends are above the target than when those are below the target. Under this regime dependent adjustment cost hypothesis, we are able to induce the model in which the changes of dividends follow the nonlinear error correction process from managers' optimization problem minimizing the costs of dividends adjustment. Using the threshold vector error correction model (VECM) proposed by Hansen and Seo (2002), we can estimate the nonlinear adjustment process of dividends. Specifically, this chapter analyzes the asymmetric adjustment behavior of the aggregate dividends in stock market with the threshold VECM, which allows for nonlinear adjustment cost and cointegration relationship. Dividends are corresponding of S&P 500 Stock Price Index. Real stock prices are used for a proxy for the target.

CHAPTER II

RATIONAL EXPECTATIONS, LONG-RUN TAYLOR RULE, AND FORECASTING INFLATION

2.1 Introduction

The rational expectations model implies that asset prices reflect forward-looking behavior in the financial market, and therefore they have been used as predictors of economic growth, business cycles, and future changes in inflation. In particular, the term structure of interest rates provides potential information on the prediction of interest rates and inflation according to the expectations hypothesis and the Fisher equation. However, the monetary authority manipulates the short-term interest rate in response to macro fundamentals such as the changes in the price level and real economic activity, and accordingly the prediction of inflation hinges on the monetary policy rules. This paper investigates the linkage between the monetary policy rules and the prediction of inflation, and provides an assessment of the predictive performance of the term structure and the monetary policy rules regarding future changes in inflation.

The predictive information contained in the yield curve has been analyzed in many empirical studies. The empirical results show that the prediction performance of the term structure model varies depending on the maturities of the yield curve and the sample period. Mishkin (1990) has shown that the term structure provides almost no information about the future change in inflation for maturities of six months or less. Fama (1990) has pointed out the variation in the real term structure, which affects the prediction performance of the term structure model. Mishkin (1991) also provided empirical results showing that the term structure provides information of predicting inflation in two or three countries out of the 10 OECD countries examined.

It is natural to ask what affects the prediction performance of the term structure model. It is noteworthy that the term structure of interest rates reveals the stylized facts of temporal persistence as discussed in Seo (2003) compared to the variation of the change in inflation. The stylized facts indicate imbalance in regression.

Many studies have shown that the persistence of the term spread is related to the monetary policy. Mankiw and Miron (1986) provided empirical results that the predictive information of the term structure began to disappear after the founding of the Federal Reserve and its manipulation of interest rates. Woodford (1999) and Rudebusch (2002) suggested that the central bank tends to adjust the target interest rate gradually, and thus such inertial monetary policy also implies the slow adjustment of the term spread. Clarida et al. (2000) have shown that the macroeconomic stability is closely related to the monetary policy rules, which involve the manipulation of the short-term interest rate as instrument to achieve the target inflation and the desirable output level. Necessarily, the variation in inflation is associated with the Fed's reaction to inflation.

Although there is a vast literature on the monetary policy rules, there have been no attempts to relate the monetary policy rules to the prediction of inflation. This paper is to provide an empirical assessment of the linkage between the monetary policy rules and the prediction of inflation. As the rational expectations model does not consider the effect of the monetary policy rules, this study resolves the mismatch between economic theory and empirical findings.

Another important issue in forecasting inflation is associated with parameter instability. The Phillips curve relates the unemployment rate to a measure of inflation. Thus, the Phillips-curve-based inflation forecasts have been used widely in monetary policymaking. However, these forecasts have been found to be sensitive to instability, particularly in the 1970s and early 1980s. Consequently, Atkeson and Ohanian (2001) argue that the likelihood of drawing an accurate prediction of a change in inflation is no better than a coin flipping. In this paper, we consider the inflation forecasts using the monetary policy rules. As the monetary policy rules may differ between the monetary policy regimes, we examine the parameter stability by using the statistical methods.

In the paper, we undertake an empirical analysis of this linkage using the U.S. monthly data for the period January 1960-December 2004. First, we estimate the long-run Taylor rule, which is composed of the federal funds rate and the 12-month inflation rate. The coefficient of reaction to inflation varies depending on the sample period and across the monetary policy regimes. Second, the prediction of inflation is found to be associated with the Fed's reaction to inflation. The coefficient of the term structure is significant for the sample period when the coefficient of reaction to inflation is close to unity. As the parameter of reaction to inflation increases, the predictive information contained in the term structure becomes weaker. This result explains the previous empirical findings that the predictive information of the term structure varies depending on the sample period. Third, an assessment of the prediction performance regarding future change in inflation is provided using the long-run Taylor rule. The empirical results indicate that the long-run Taylor rule improves forecasting accuracy.

The paper is organized as follows. Section 2.2 deals with the term structure model and the influence of the monetary policy rules on the prediction of inflation. Section 2.3 discusses the econometric methods to assess the information contained in the term structure and the long-run Taylor rule. The main results are provided in Section 2.4.

2.2 The Model

The Fisher equation implies that the nominal interest rates reflect expectations of inflation, and therefore the term structure provides potential and useful information about the future path of inflation. Fama (1990) and Mishkin (1990) assessed the predictive information contained in the term structure based on the following model.

$$\pi_{t,t+m} - \pi_{t,t+l} = \mu + \alpha(R_t^m - R_t^l) + u_{t+m}, \tag{2.1}$$

where $\pi_{t,t+h}$ is the h-step ahead inflation, and R_t^h is the nominal yield on a security with a maturity of h for $h=m$, l and $m>l$.

The term structure model (2.1) implies that the change in inflation depends on the term structure of interest rates. From the Fisher equation, the nominal interest rate (R_t^h) is composed of the real interest rate (κ_t^h) and the expected inflation as follows:

$$R_t^h = \kappa_t^h + E_t(\pi_{t,t+h}), \tag{2.2}$$

where $E_t(\cdot)$ is the conditional expectation based on the information available at time t.

By taking a difference of l-step ahead inflation from m-step ahead inflation, we get the term structure model (2.1) and the following conditions.

$$\mu = -E(\kappa_t^m - \kappa_t^l)$$

$$\alpha = 1$$

$$u_{t+m} = [\pi_{t,t+m} - E_t(\pi_{t,t+m})] - [\pi_{t,t+l} - E_t(\pi_{t,t+l})] - [(\kappa_t^m - \kappa_t^l) - E_t(\kappa_t^m - \kappa_t^l)].$$

If we assume rational expectations and the constancy of the real term structure, $E_t(u_{t+m}) = 0$ holds in equation (2.1) and the error u_{t+m} is exogenous to the variables in the current information set. As a result, the future change in inflation has a linear relationship with the term structure with a unit slope. Therefore, the term structure provides systematic information about the future path of inflation.

The prediction performance of the term structure model has been examined in many empirical studies. The results show that the predictability of inflation varies depending on the maturities of the yield curve and the sample period. Mishkin (1990) has shown that the term structure of interest rates provides almost no information about the future change in inflation for maturities of six months or less. Fama (1990) has pointed out the variation in the real term structure, which brings in less-successful performance of the term structure model. Mishkin (1991) also provided empirical results showing that the term structure provides information of predicting inflation in two or three countries out of the 10 OECD countries examined.

It is natural to ask what affects the prediction of inflation based on the rational expectations model. One plausible explanation, suggested in previous studies, is related to the non-spherical errors, which may affect the prediction performance of the term structure model. The term structure model involves the overlapping data, which generates serial correlation in the error term inevitably. However, the problem of overlapping data becomes more severe for long-period ahead inflation forecasting while the empirical evidences are less favorable in forecasting inflation for maturities of six months or less.

The term structure of interest rates reveals the stylized facts of temporal persistence and nonlinear mean reversion as shown by Seo (2003). On the other hand, the change in inflation is relatively less persistent, and thus the stylized facts indicate imbalance between the term structure and the change in inflation.

It has been shown in many studies that the persistence of the term spread is related to the monetary policy. Mankiw and Miron (1986) provided empirical results that the predictability of the term structure began to disappear after the founding of the Federal Reserve and its manipulation of interest rates. Rudebusch (1995) and Balduzzi et al. (1997) also found that the changes in the interest rate were due to the Fed's unexpected changes in its target interest rate. As Woodford (1999) suggests, the central bank tends to adjust interest rates gradually, and thus such inertial monetary policy also implies the slow adjustment of the term spread.

According to the expectations hypothesis, the long-term interest rate is the average of the current and future short-term interest rates.

$$R_t^m = \frac{1}{m}\sum_{i=1}^{m} E_t(R_{t+i-1}) + q_t, \tag{2.3}$$

where R_t^m is the yield on a security with a maturity of *m*, R_t is the yield on the unit-maturity security, and q_t is the liquidity premium.

The expectations hypothesis (2.3) can be written as follows:

$$R_t^m - R_t = \frac{1}{m}\sum_{i=1}^{m-1}\sum_{j=i}^{m-1} E_t(\Delta R_{t+m-j}) + q_t.$$

If the liquidity premium is constant, the expectations hypothesis implies that the term structure or the yield curve provides information on the future change in the short-term interest rate. Thus, the expectations hypothesis implies that the change in the short-term interest rate depends on the term structure. However, the empirical findings suggest that the persistence of the term structure is closely related to the Fed's control of interest rates. In particular, Taylor (1993) suggested the monetary policy rules. The monetary authority regulates the target interest rate (r_t^*) in response to the macro fundamentals: one-year inflation rate (π_t) and output gap (y_t) as follows.

$$r_t^* = r^* + \beta(\pi_t - \pi^*) + \theta y_t, \tag{2.4}$$

where r^* is the desired nominal rate, which is compatible with the inflation target π^*.

The Fed's reaction function has been estimated by assuming the partial adjustment process in Clarida et al. (2000) and Rudebusch (2002).

$$\begin{aligned} r_t &= (1-\rho)r_t^* + \rho r_{t-1} \\ &= (1-\rho)(\beta\pi_t + \theta y_t + v) + \rho r_{t-1}, \end{aligned}$$

where r_t is the actual federal funds rate and $v = r^* - \beta\pi^*$. Rudebusch (2002) estimated the reaction function and found that the partial adjustment coefficient ρ is large and significant, which supports the monetary policy inertia. Judd and Rudebusch (1998) used the error correction specification because the unit root hypotheses of the interest rates cannot be rejected.

$$\begin{aligned}\Delta r_{t+1} &= \phi(r_t - r_t^*) + C(L)\Delta r_t \\ &= \phi(r_t - \beta\pi_t - \theta y_t - v) + C(L)\Delta r_t.\end{aligned}$$

If ϕ<0, the federal funds rate adjusts to the equilibrium error between the actual funds rate and the optimal target rate. The equilibrium error disappears eventually, which implies a long-run equilibrium relationship. The long-run relationship is governed by two highly persistent variables: the federal funds rate and the inflation rate.

$$w_t = r_t - \beta\pi_t. \tag{2.5}$$

The long-run coefficient β is the parameter of reaction to inflation. If w_t is stationary, the long-run monetary policy rules form a long-run relationship based on the definition of Engle and Granger (1987). The output gap is stationary, and it affects the long-run relationship temporarily. This specification makes our empirical analysis simple and tractable. However, our analysis can be extended to the monetary policy rules that include real economic activity. If we include the output gap, the influence of the monetary policy rules on the prediction of inflation can be explained by the variation in the output gap.

The rational expectations model does not consider the Fed's control of interest rates in response to inflation. The expectations hypothesis implies the long-run relationship between the short rate and the long rate. However, if the monetary policy rules are effective, the short rate converges to the target rate, which can be different from the long rate. Thus, the relationship between the term structure and the change in inflation becomes weaker.

The long-run relationship w_t can be written as follows:

$$w_t = (r_t - R_t) + (R_t - \beta E_t \pi_{t+m}) + \beta (E_t \pi_{t+m} - \pi_t).$$

The long-run Taylor rule w_t is composed of the term spread, the relationship between the long-term rate and the expected inflation, and the expected change in inflation. Accordingly, the long-run monetary policy rules imply a relationship between the term structure and the change in inflation.

$$\pi_{t+m} - \pi_t = \frac{1}{\beta}(R_t - r_t) + \eta_{t+m}, \tag{2.6}$$

where

$$\eta_{t+m} = \frac{1}{\beta}[(r_t - \beta \pi_t) - (R_t - \beta E_t \pi_{t+m})] + (\pi_{t+m} - E_t \pi_{t+m}).$$

If the long-run parameter β equals one, the long-run Taylor rule reduces to the short-term realized real interest rate. Also, the relationship between the long-term rate and the expected inflation becomes the long-term real interest rate. If we assume the constancy of the real term structure, the implied term structure model becomes close to the rational expectations model. In that case, the long-run monetary policy rules are consistent with the rational expectations model.

However, this is a special case. If β is different from one, the slope and the error in (2.6) depend on the parameter β. First, an increase in the long-run reaction parameter leads to a decrease in the slope, which lowers the effect of the term structure in predicting inflation. Second, if β is different from one, the term structure model is valid under the constancy of the long-run monetary policy rules. In general, the change in inflation depends on the long-run monetary policy rules as well as the term structure. Third, the discrepancy between the Fisher equation and the long-run monetary policy rules increases as β increases. The discrepancy generates uncertainty in forecasting inflation, and as a result the variance of the error increases and the relevancy of the forecasts may diminish. Finally, the prediction performance of the term structure model can be affected by parameter uncertainty in the reaction parameter β.

The parameter uncertainty cannot be overlooked because it affects the prediction of inflation severely. Clarida et al. (2000) related the monetary policy rules to macroeconomic stability. The reaction parameter may change across the monetary policy regimes, which generates parameter uncertainty in forecasting inflation. Furthermore, the Fed's reaction may vary over the business cycle. The monetary authority is likely to focus on the prevention of inflation in the boom while high unemployment becomes the main concern in the recession. The central bank's regime-dependent preferences have been suggested in Ruge-Murcia (2003), which also produces parameter uncertainty in forecasting inflation.

When the long-run monetary policy rules include other macro fundamentals, uncertainty in forecasting inflation inevitably increases. In addition, the term structure is associated with real economic activity as shown by Estrella and Hardouvelis (1991), and the measurement of output gap accompanies informational limitation as discussed in

Orphanides (2003). These factors increase uncertainty and reduce the relevancy of the inflation forecasts.

The predictability of the term structure model has been measured in many studies. However, the assessment of the term structure information has been based on the rational expectations model, and the long-run aspects of the monetary policy rules have not been considered. In this study, we examine the prediction of inflation using the long-run information contained in the monetary policy rules.

2.3 Methodology

2.3.1 Forecasting Models

Denote π_t as the 12-month inflation rate, r_t as the federal funds rate, and R_t as the yield on the one-year Treasury note. Our model of forecasting inflation is based on the following:

$$\pi_{t+m} - \pi_t = \mu + \alpha(R_t - r_t) + \lambda(r_t - \beta\pi_t) + \sum_{i=1}^{k} \gamma_i \Delta\pi_{t-i} + \eta_{t+m}. \tag{2.7}$$

Our model (2.7) is very close to the forecasting model used by Stock and Watson (1999), which explains the change in inflation using the term structure information. Our forecasting model incorporates the information of the long-run monetary policy rules. The long-run Taylor rule accompanies the parameter β, which signifies the Fed's reaction to inflation. In the paper, we estimate the long-run parameter β by using reduced rank regression on the vector error correction model. The lagged values of the differenced inflation are added to reduce serial correlation in the error. If $\lambda = 0$, our model becomes the term structure model as follows:

$$\pi_{t+m} - \pi_t = \mu + \alpha(R_t - r_t) + \sum_{i=1}^{k} \gamma_i \Delta\pi_{t-i} + \eta_{t+m}. \tag{2.8}$$

Thus, if the long-run information of the monetary policy rules does not help explain the change in inflation, our model reduces to the forecasting model using the term structure information, which has been proposed by Stock and Watson (1999).

The Martingale property of inflation has been suggested in several studies such as Atkeson and Ohanian (2001). The Martingale property implies that the future change in inflation is unpredictable. We treat the random walk model as the reference model to evaluate the inflation forecasting models.

$$\pi_{t+m} - \pi_t = \mu + \eta_{t+m}. \tag{2.9}$$

We compare the predictive performance of the inflation forecasting models-Model A: the random walk model; Model B: the forecasting model that uses the term structure; and Model C: the forecasting model that uses the term structure and the long-run monetary policy rules.

2.3.2 Parameter Stability

When we evaluate the forecasting models, we need to consider parameter uncertainty because it affects the prediction accuracy severely. As discussed in Clarida et al. (2000), the monetary policy rules may differ between the monetary policy regimes. To examine the parameter stability, we implement the tests for structural change in the reaction parameter of the Taylor rule.

$$r_t = \beta_1 \pi_t 1(t \le t^*) + \beta_2 \pi_t 1(t > t^*) + \omega_{2t}, \tag{2.10}$$

where $1(\cdot)$ is the indicator function, and t^* is the date of the break point.

In policy regime 1, the Fed reacts to inflation by adjusting the target rate with the coefficient β_1. In policy regime 2, the magnitude of reaction may change depending on the coefficient β_2. If the magnitude of reaction to inflation does not vary across regimes, the linear error correction model is valid. Therefore, the tests for structural change in the long-run Taylor rule can be based on the following hypotheses:

$$H_0 : \beta_1 = \beta_2 \text{ against } H_1 : \beta_1 \ne \beta_2.$$

We assume that the date of structural change is unknown. Although the dates of the monetary policy regimes are known, it is the general case that the true date of break may differ from the historical date. Thus, the testing for structural change entails the nuisance parameter t^*, which cannot be identified under the null hypothesis as discussed in Andrews (1993). We use the optimal test statistics defined in Seo (1998).

$$AveLM_n = \frac{1}{t_U - t_L + 1} \sum_{t^* = t_L}^{t_U} LM_n(t^*),$$

$$ExpLM_n = \log[\frac{1}{t_U - t_L + 1} \sum_{t^* = t_L}^{t_U} \exp(LM_n(t^*)/2)],$$

$$SupLM_n = Max_{t^* \in [t_L, t_U]} LM_n(t^*).$$

The algorithm to compute the test statistics is as follows. First, we estimate the linear error correction model. Second, we calculate the LM statistics using the null model and parameter estimates for each break point $t^* \in [t_L, t_U]$. The trimming values can be chosen symmetrically with the trimming probability P, for example, .10 or .15. Third, we find the average, the weighted average, and the maximum of the LM statistics. As the test statistics follow nonstandard distributions, we use the critical values suggested in Seo (1998). If the test statistic is greater than the critical value, we reject the null hypothesis of no structural change.

2.4 Main Results

In the empirical analysis, we use the monthly data of the federal funds rate ($=r_t$) and the yield on the one-year U.S. Treasury note ($=R_t$). The 12-month inflation rate is calculated using the consumer price index (CPI). That is, $\pi_t = (\log P_t - \log P_{t-12}) \times 100$, where P_t is the CPI.

The data set is obtained from the Federal Reserve Economic Data[1] for the sample period January 1960-December 2004 (1960:1-2004:12). The estimation of the model and the in-sample forecasts are based on the sample period 1960:1-1999:12. The out-of-sample forecasts are obtained for the period 2000:1-2004:12.

[1] Http://research.stlouisfed.org/fred2.

Figure 2.1 shows the change in inflation of 12-month horizon, which is $\pi_{t+12} - \pi_t$. The time plot of the term spread is provided in Figure 2.2. The term spread, defined as $R_t - r_t$, varies slowly compared to the variation of the change in inflation.

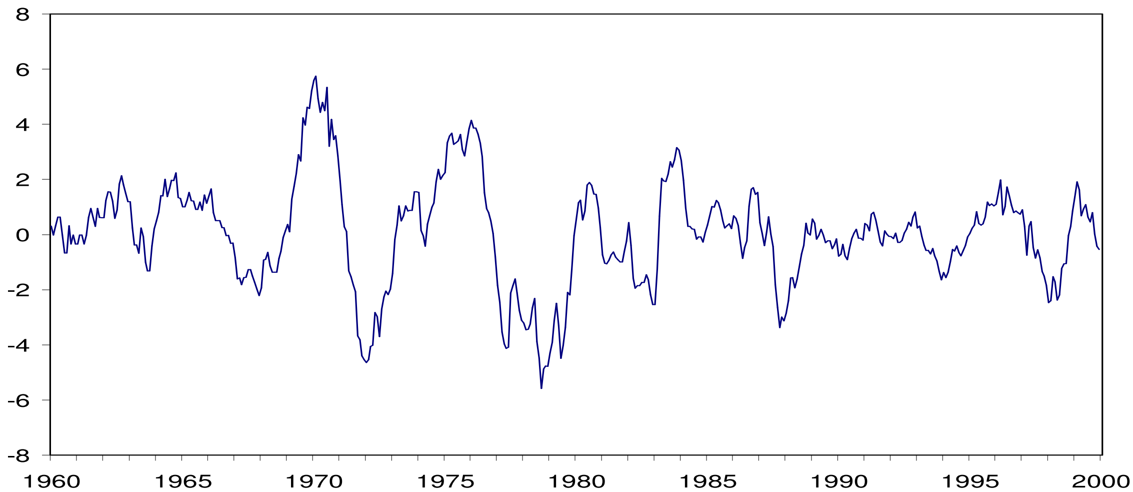


Figure 2.1 Change in Inflation

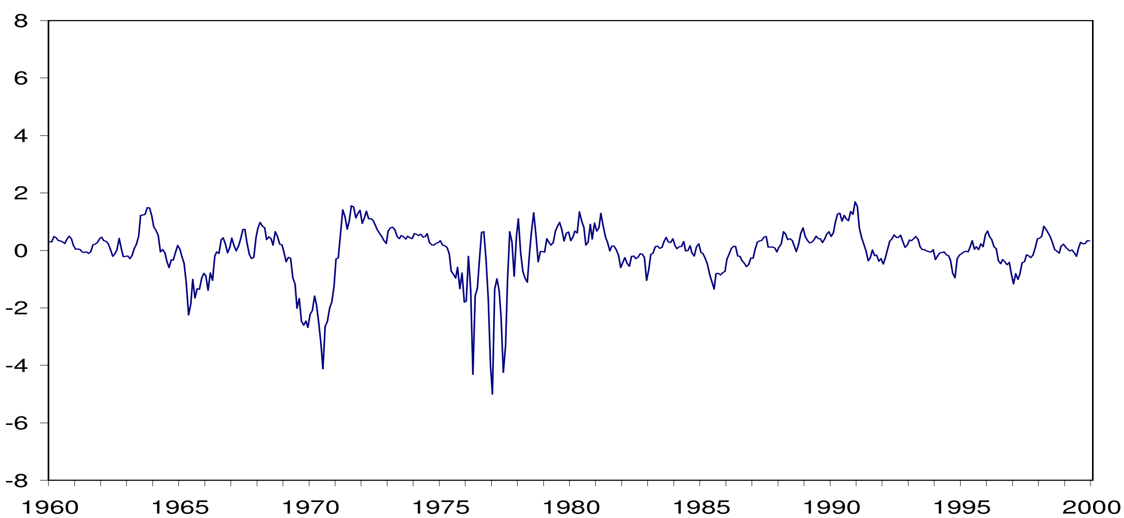


Figure 2.2 Term Spread

Because the term structure predictability may depend on the monetary policy rules, we investigate this linkage statistically. Our empirical analysis involves the estimation of the long-run Taylor rule, and so we examine the time series behavior of the variables to estimate the long-run Taylor rule. Table 2.1 shows the augmented Dickey-Fuller (ADF) unit root tests. The unit root hypothesis of the 12-month inflation rate cannot be rejected for each AR lag length from 1 to 7. The federal funds rate shows mixed results. At the AR lag length 2, the ADF test rejects the unit root hypothesis while the unit root hypothesis maintains at other lag lengths. At the AR lag length 3, which is chosen by the Bayesian information criterion (BIC), the ADF test cannot reject the null hypothesis of unit root in the federal funds rate. The yield on the one-year Treasury note is persistent and the unit root hypotheses cannot be rejected.

Table 2.1
Unit Root Tests

Variables	AR Lag Length						
	1	2	3	4	5	6	7
π_t	-1.470	-1.895	-2.160	-2.046	-2.165	-2.463	-2.507
r_t	-2.315	-3.290	-2.811	-2.727	-2.501	-2.484	-2.367
R_t	-2.169	-2.984	-2.362	-2.371	-2.279	-2.523	-1.996

The critical value at the 5% significance level is -2.867.

Table 2.2 shows the cointegration tests for the term structure and the long-run Taylor rule, which is composed of the federal funds rate and the 12-month inflation rate. The long-run Taylor rule implies that these two variables have a long-run relationship. At the VAR lag order 2, the Johansen cointegration test rejects the null hypothesis of no cointegration at the 5% significance level. However, at the VAR lag order 3, which is chosen by the BIC, the trace statistic for cointegration is slightly less than the 5% critical value. The cointegration tests support the long-run relationship of the term structure between the federal funds rate and the long-term interest rate at each VAR lag length. Therefore, the term structure contains the long-run information of predicting the short-term interest rate.

Table 2.2
Cointegration Tests

Variables	VAR Lag Length					
	1	2	3	4	5	6
(r_t, π_t)	13.252	23.114	19.223	17.736	16.435	17.977
(r_t, R_t)	53.737	57.543	39.293	34.019	24.414	24.786

5% critical value is 20.262. The VAR lag length selected by the BIC is 3 for each model.

Using the bivariate error correction model, the long-run Taylor rule is estimated at the VAR lag length 3, which is chosen by the BIC. As shown in Table 2.3, for the sample period 1960:1-1999:12, the long-run coefficient is close to one, which is compatible with the rational expectations model. However, the reaction coefficient varies widely across the monetary policy regimes. The magnitude of reaction to inflation increased in the Greenspan monetary policy regime (1987:8-1999:12) compared to the entire in-sample period. The reaction coefficient is large, and its standard error is also huge, which reflects the variation in the Fed's reaction to inflation.

Table 2.3
Long-run Taylor Rule

Sample Period	β		ν	
1960:01-1999:12	0.900	(0.191)	2.701	(0.997)
1987:08-1999:12	2.629	(0.861)	-2.464	(2.934)

Estimation model is as follows: $r_t = \nu + \beta\pi_t + \omega_t$. The standard errors are in the parentheses.

Our model implies that the term structure information loses its predictability of inflation as the magnitude of reaction to inflation increases. At the same time, the parameter uncertainty is likely to lower the relevancy of the inflation forecasts based on the term structure information.

Table 2.4 shows the results of testing for parameter stability of the long-run Taylor rule. The test statistics are based on the bivariate error correction model of the federal

Table 2.4
Parameter Stability of the Long-run Taylor Rule

	Ave-LM	5% c.v.	Exp-LM	5% c.v.	Sup-LM	5% c.v.
β	1.086	2.71	3.46	2.02	17.283	9.09
$adj.\ vector$	7.343	4.61	10.824	3.22	28.761	11.79
$(\beta, adj.\ vector)$	8.429	6.08	13.524	4.25	36.247	14.23

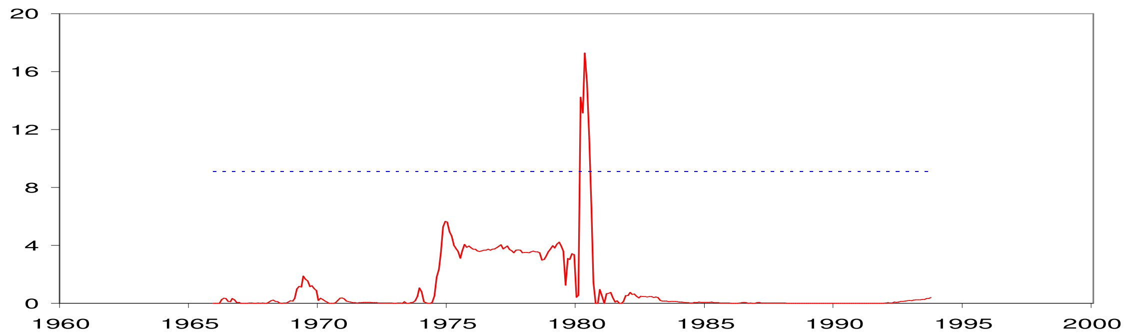
The 5% critical values are in the parentheses.

funds rate and the 12-month inflation rate for the sample period 1960:1-1999:12. The 5% critical values are obtained from Seo (1998) for the stability of the long-run cointegrating vector and from Andrews (1993) for the stability of the adjustment vector.

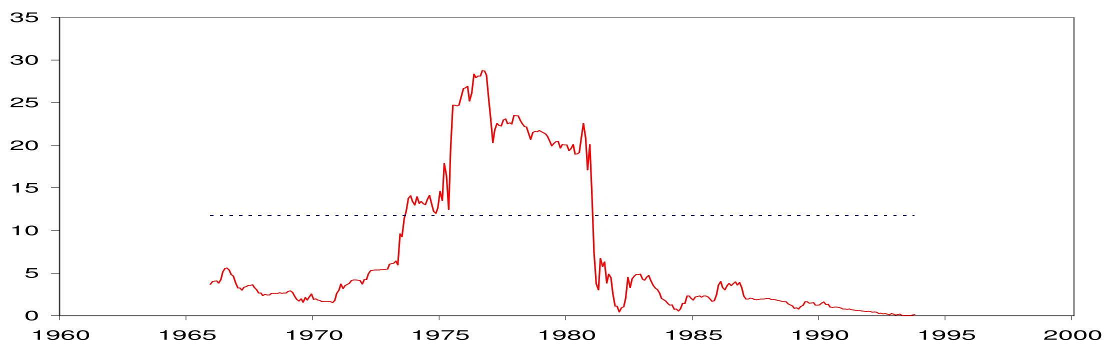
The parameter stability of the long-run reaction parameter can be rejected based on the Exp-LM and Sup-LM statistics. Although the Ave-LM statistic does not support parameter instability, Figure 2.3 shows that parameter instability increased in the mid 1970s and reached the peak in the early 1980s. This result coincides with the period of the change in the operating system for which the volatility of the interest rate and inflation increased. After the mid 1980s, the LM statistics of the long-run reaction parameter became stabilized. Also, the parameter stability of the short-run adjustment vector can be rejected. We find parameter instability in the Fed's reaction to inflation. Parameter uncertainty may affect the relevancy of the inflation forecasts.

Next, we compare the prediction accuracy of inflation forecasting models: random walk; forecasting with the term structure; and forecasting with the long-run Taylor rule and the term structure.

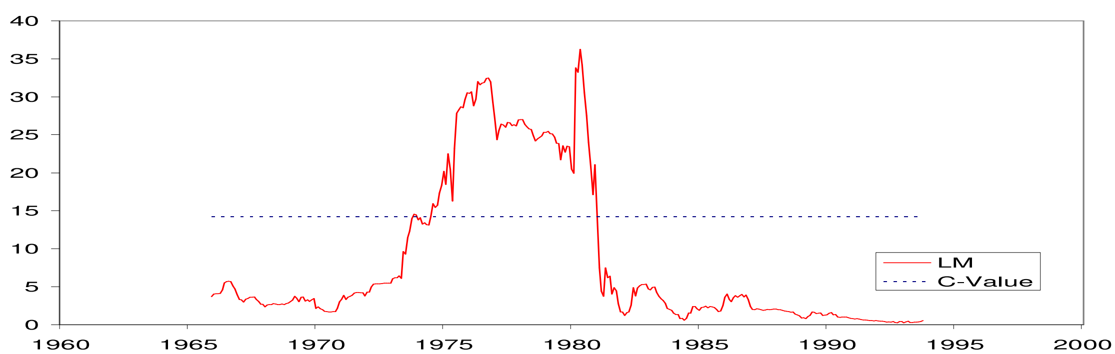
Table 2.5 reports estimation results of the forecasting models. First, we estimate the forecasting model using the term structure. An intercept and four lagged values ($k=4$) of differenced inflation are augmented to estimate the forecasting model. For the sample period 1960:1-1999:12, the response of inflation to the term structure is significant although the term spread has the limited predictability of the change of inflation as the adjusted R-squared coefficient shows. However, for the period 1987:8-1999:12, the response of inflation to the term structure became negative and insignificant.



Panel A: Long-run reaction parameter



Panel B: Adjustment coefficient



Panel C: Joint test of long-run reaction parameter and adjustment coefficient

Figure 2.3 Stability Tests of the Long-run Taylor Rule

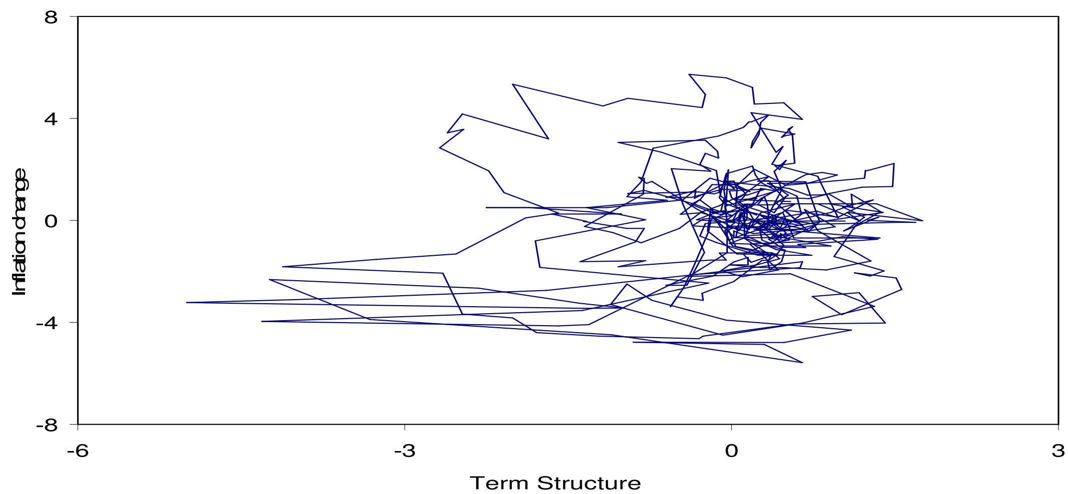
Table 2.5
Inflation Forecasting Model

	$R_t - r_t$		$r_t - \beta\pi_t$		$\bar{R}^2$
1960:1-1999:12	0.393	(0.183)			0.078
	0.426	(0.226)	0.035	(0.117)	0.077
1987:8-1999:12	-0.187	(0.253)			0.009
	-0.075	(0.233)	0.255	(0.071)	0.343

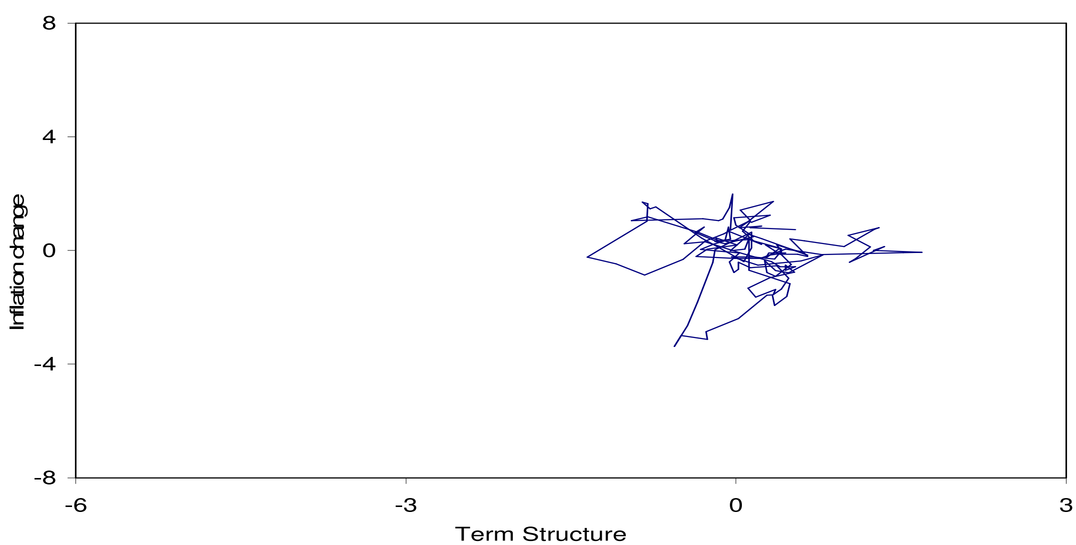
The standard errors are in the parentheses.

Figure 2.4 depicts the relationship between inflation change and term spread, which supports the estimation results. As Figure 2.4 shows, the change in inflation is weakly related to the term spread for the entire in-sample period. However, this relationship disappeared in the Greenspan monetary policy regime.

For the sample period 1960:1-1999:12, the long-run information of the Taylor rule is not significant as shown in Table 2.5. However, for the sample period 1987:8-1999:12, the predictability of the model with the long-run Taylor rule improves dramatically in terms of the adjusted *R*-squared coefficient compared to the forecasting model using the term spread only. The information of the long-run Taylor rule is calculated using the estimated reaction parameter. In addition, an intercept and four lagged values ($k=4$) of differenced inflation are augmented to estimate the model. While the term structure information is weak in the Greenspan monetary policy regime, the long-run Taylor rule exhibits a significant information effects. The change in inflation responds positively to the long-run Taylor rule. When the actual short-term rate is greater than the optimal target rate, the equilibrium process begins with an increase in inflation. Therefore, the long-run Taylor rule provides information to predict the future change in inflation. Figure 2.5 displays the relationship between the change in inflation and the long-run Taylor rule. This relationship becomes evident for the Greenspan monetary policy regime.

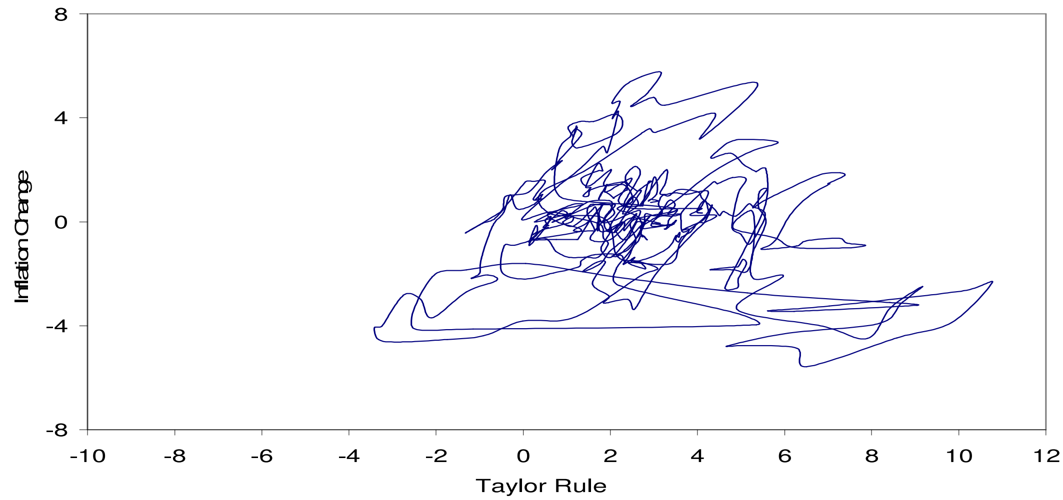


Panel A: 1960:1-1999:12

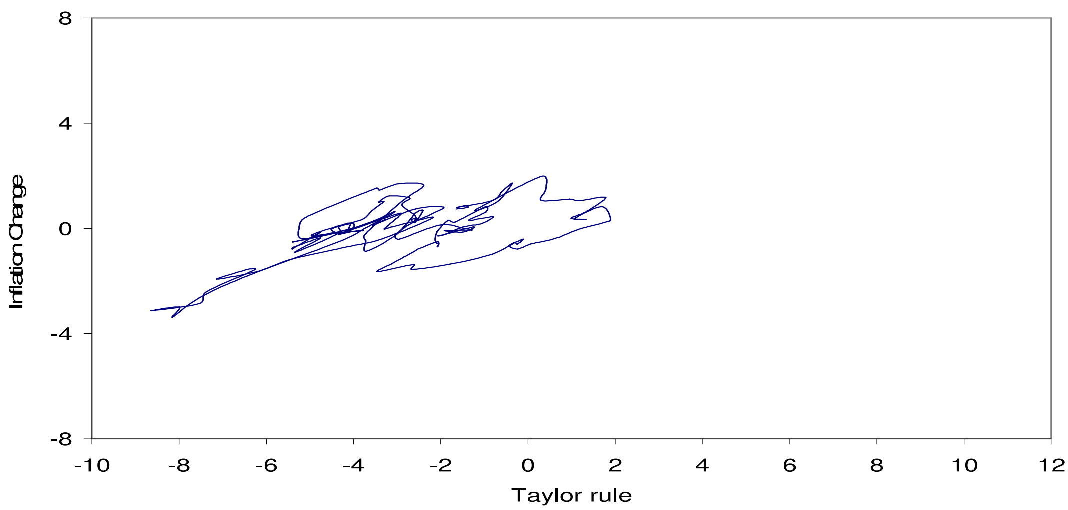


Panel B: 1987:8-1999:12

Figure 2.4 Term Structure and Inflation Change



Panel A: 1960:1-1999:12



Panel B: 1987:8-1999:12

Figure 2.5 Taylor Rule and Inflation Change

We examine the robustness of the predictive information in the long-run Taylor rule by using the different forms of inflation forecasting model. First, we consider the term structure of interest rates with different maturities. The term spread is defined as the difference of the yields between the 10-year Treasury bond and 3-month Treasury bill. As Table 2.6 shows, the coefficient of the term spread has the negative sign and it is insignificant for the period 1960:1-1999:12. However, the long-run Taylor rule has a significant information effect in predicting the change in inflation for the Greenspan monetary policy regime. The similar results, which are not reported in the paper, are obtained for several choices of the term structure with different maturities.

Table 2.6
Inflation Forecasting Model with Different Term Structure

	$R_t - r_t$		$r_t - \beta\pi_t$		$\overline{R}^2$
1960:1-1999:12	-0.066	(0.142)			0.046
	-0.084	(0.139)	-0.044	(0.112)	0.046
1987:8-1999:12	-0.207	(0.095)			0.052
	0.113	(0.157)	0.286	(0.088)	0.353

The standard errors are in the parentheses.

We also consider several macroeconomic variables in inflation forecasting model with the long-run Taylor rule and the term structure. As Table 2.7 shows, the coefficients of term structure and the Taylor rule do not appear to be seriously affected by the inclusion of macroeconomic variables. For the sample period 1960:1-1999:12, the coefficients of the macroeconomic variables such as unemployment rate, the change in industrial production, M2 growth, and the oil price change are significant in explaining the change in inflation. However, these macroeconomic variables become insignificant for the sample period 1987:8-1999:12.

Table 2.7
Inflation Forecasting Model with Other Macro Variables

Sample Period	1960:01-1999:12		1987:8-1999:12	
Term Structure	0.3646	(0.1686)	-0.1757	(0.2498)
Taylor Rule	0.0750	(0.0869)	0.3001	(0.0923)
Unemployment	-0.4118	(0.1003)	0.1389	(0.1629)
IP Change	0.0536	(0.0124)	0.0022	(0.0126)
M2 Growth	0.0745	(0.0315)	-0.0043	(0.0316)
S&P 500 Returns	0.0026	(0.0020)	-0.0014	(0.0015)
Oil Price Change	0.0021	(0.0008)	0.0013	(0.0009)
$\overline{R}^2$	0.270		0.347	

The standard errors are in the parentheses.

Table 2.8 summarizes the predictive accuracy of inflation forecasting models. The random walk model is treated as the reference model. The inflation forecasts using the long-run Taylor rule and the term structure achieve an improvement in the predictive accuracy by 4.55% in terms of the RMSE compared to the random walk model for the sample period 1960:1-1999:12. The MAE decreases by 2.02% for the same period. On the other hand, for the sample period 1987:8-1999:12, the inflation forecasts using the long-run Taylor rule show an improvement in the prediction accuracy by 20.62% in terms of the RMSE relative to the random walk model while the term structure information reveals 2.14% gain. Therefore, the inflation forecasts using long-run Taylor rule information show an improvement in the prediction accuracy relative to the forecasts using the term structure only.

Table 2.8
Forecasting Accuracy

		Model 1 (=A)	Model 2 (=B)	Model 3 (=C)	B/A	C/A
		In-Sample forecasting				
1960:1-1999:12	RMSE	1.8746	1.7906	1.7893	0.9552	0.9545
	MAE	1.3689	1.3430	1.3412	0.9811	0.9798
1987:8-1999:12	RMSE	0.9609	0.9403	0.7627	0.9786	0.7938
	MAE	0.7029	0.7034	0.6199	1.0007	0.8819
		Out-of-Sample forecasting				
1960:1-1999:12	RMSE	1.2036	1.2704	1.2803	1.0554	1.0637
	MAE	1.0181	1.0977	1.1133	1.0782	1.0934
1987:8-1999:12	RMSE	1.1914	1.1295	1.0568	0.9481	0.8871
	MAE	1.0307	0.9640	0.9059	0.9353	0.8790

$\text{RMSE} = \sqrt{\frac{1}{n}\sum_{t=1}^{n}(\pi_t - \hat{\pi}_t)^2}$; $\text{MAE} = \frac{1}{n}\sum_{t=1}^{n}|\pi_t - \hat{\pi}_t|$

Table 2.8 also shows the prediction accuracy of the out-of-sample forecasts for the period 2000:1-2004:12. The forecasts are calculated recursively with a start-up sample period of 1960:1-1999:12 and 1987:8-1999:12. Given the start-up sample period 1960:1-1999:12, the out-of-sample inflation forecasts do not show any improvement regardless of the information about the term structure and the long-run Taylor rule. However, given the start-up sample period 1987:8-1999:12, the out-of-sample inflation forecasts using the long-run Taylor rule and the term structure achieve a significant improvement in the predictive accuracy by 11.29% measured by the RMSE while the out-of-sample forecasts using the term structure only improves 5.19% compared to the random walk model. Considering parameter instability in the monetary policy rules, this evidence is quite noteworthy. As the parameter in the monetary policy rules becomes more stable, the inflation forecasts using the long-run Taylor rule are likely to generate more accurate prediction of inflation.

2.5 Conclusion

In this paper, we investigate the influence of the monetary policy rules on the prediction of inflation. Our analysis finds that the prediction performance of the term structure model hinges on the monetary policy rules, which involve the manipulation of the federal funds rate in response to the change in the price level. As the Fed's reaction to inflation becomes stronger, the predictive information contained in the term structure becomes weaker. Using the long-run Taylor rule, a new assessment of the prediction performance regarding future change in inflation is provided. The empirical results indicate that the long-run Taylor rule improves forecasting accuracy. The rational expectations model cannot explain this linkage, and thus this study resolves the discordance between economic theory and empirical findings.

We extended our analysis to the model with other macroeconomic variables. The information of economic indicators tends to be less important as the central bank shows strong commitment to the inflation. However, the information of the monetary policy rules, if strong, can be used for predicting the future path of inflation.

CHAPTER III

IS THE PREFERENCE OF CENTRAL BANK ASYMMETRY? EVIDENCE FROM KOREA

3.1 Introduction

There are a number of studies that have investigated the optimal monetary policy or the monetary policy rules under the assumption that the preferences of the central banks are symmetric over key macroeconomic variables, such as inflation and output. Usually they employ the standard linear-quadratic framework which consists of central banks' quadratic loss function and a linear aggregate supply function and/or a linear demand function. This quadratic (symmetric) preference means that the central banks have weighted the same amount of loss both on positive deviations of inflation (output) from target (trend) and on negative ones of the same size. This linear-quadratic framework leads to a linear reaction function or Taylor rule type function which is the first order condition of the optimization of central banks' problem.

However, recently a growing number of papers have questioned this quadratic preference assumption: specifically, are negative deviations of inflation from target as undesirable as positive deviations of inflation from target of the same amount? And/or are positive output gaps as distasteful as negative output gaps of the same size? Nobay and Peel (2003), Ruge-Murcia (2002), Surcio (2003a), and Karagedikli and Lees (2004) relax the assumption of the quadratic preference of central banks and adopt instead asymmetric preference specifications. This asymmetric loss function generally leads to a nonlinear reaction function or monetary policy rule which is the first order condition solved for the optimization problem of central banks with a linear aggregate supply function. Whether the central bank has the asymmetric objective function is an important issue since many of the results on the time consistency problem under symmetric preferences may no longer hold under asymmetric preferences as shown in Nobay and Peel (2003), and Surcio (2003a).

In principle the preference of central banks can be inferred from their monetary policy reaction function. However, a nonlinear monetary policy reaction function can also be derived from a nonlinear aggregate supply curve. Therefore, Surcio (2003b), Dolado et al. (2004) studied central banks' asymmetric preference with a nonlinear aggregate supply curve. They tried to estimate the asymmetric preferences of central banks controlling the nonlinear component coming from the nonlinear aggregate supply curve.

Recently the way of conducting monetary policy in Korea was changed dramatically. After the Bank of Korea Act revised in April 1998, the inflation targeting was adopted as new monetary policy system. In addition the Bank of Korea (i.e., BOK), which is the central bank in Korea, uses interest rate - the overnight call rate - as its monetary policy instrument instead of monetary aggregates such as reserves, and M2. Table 3.1 shows the change of the target interest rate in the period from 1999 to 2005. This institutional change has stimulated Taylor rule type monetary policy analysis of the Bank of Korea. For example, Eichengreen (2004) analyzed the monetary and exchange rate policy of the

Bank of Korea by estimating the Taylor type monetary policy rule. Since the Bank of Korea has set up the explicit inflation target from 1998, the study on its asymmetric preference of inflation will have an advantage over assuming the implicit inflation target frequently used for the non-inflation targeting countries, for example United States.

Table 3.1
The Change of the Target Interest Rates

Date of change		The target rate (%)	The change (%)
1999	May 6	4.75	-
2000	February 10	5.00	+0.25
	October 5	5.25	+0.25
2001	February 8	5.00	-0.25
	July 5	4.75	-0.25
	August 9	4.50	-0.25
	September 19	4.00	-0.25
2002	May 7	4.25	+0.25
2003	May 13	4.00	-0.25
	July 10	3.75	-0.25
2004	August 12	3.50	-0.25
	November 11	3.25	-0.25
2005	October 11	3.50	+0.25
	December 8	3.75	+0.25

Sources are the Bank of Korea.

Most of studies about the asymmetric preference of central banks are concentrated on the developed countries. Also empirical results are not conclusive on asymmetric preference of central banks so that they are dependent on the countries and the sample period studied. However, to the best of my knowledge still there is no study of the asymmetric preference of the central bank of Korea.

This chapter investigates the monetary policy preference of the central bank in Korea during the inflation targeting period: 1998:9-2005:12. We are trying to answer the question whether the preferences for inflation or/and output gap of the central bank in Korea are asymmetry, and specifically the BOK has a precautionary demand for inflation or/and for expansions.[2] The contribution of this paper is to provide some evidence supporting that the preference of central banks may not be symmetric over inflation but over output gap, and therefore it adds another empirical result to the literature of asymmetric preferences of central banks.

The rest of this chapter is organized as follows. Section 3.2 presents the new Keynesian economy and the asymmetric objective function of central bank along line of Surcio (2003b), etc. Under this framework, we derive the optimal reaction function which serves as a benchmark for the empirical section. The empirical results are provided in Section 3.3. Section 3.4 concludes.

3.2 The Model

3.2.1 The Economy

In this chapter we adopt the new Keynesian framework as an economy in which inflation and output gap depend on the expected future values of those variables respectively and in which the policy instrument of the central bank is the nominal

[2] Cukierman and Muscatelli (2003) call it the situation that when policy makers are uncertain the state of the economy, they might respond more aggressively to negative output gaps than to positive ones.

interest rate. This simple model has been popularized for use in monetary policy literature after Clarida et al. (1999), Woodford (1999, 2001), McCallum and Nelson (1999) and Svensson and Woodford (1997, 2003), etc. Specifically the economy can be specified by the following two equation system corresponding to an aggregate demand and to an aggregate supply relation, respectively:

$$\tilde{y}_t = E_t \tilde{y}_{t+1} - \alpha(i_t - E_t \pi_{t+1}) + e_t^d, \tag{3.1}$$

$$\pi_t = \omega E_t \pi_{t+1} + F(\tilde{y}_t) + e_t^s, \tag{3.2}$$

where

$$F(\tilde{y}_t) = \beta \tilde{y}_t / (1 - \beta \psi \tilde{y}_t). \tag{3.3}$$

$\tilde{y}_t$ is the output gap and π_t is the inflation rate, $E_t \tilde{y}_{t+1}$ and $E_t \pi_{t+1}$ are the expected values of those variables conditioned on the information available at period *t*, i_t is the nominal rate of interest, and $\alpha > 0$, $\omega > 0$, $\beta > 0$ and $\psi \geq 0$ are parameters. The demand disturbance, e_t^d, and the cost disturbance, e_t^s, are assumed to follow a zero mean reverting process. Equation (3.1) is represented by a linear approximation to the representative household's Euler condition for optimal consumption. It basically postulates a forward-looking IS relationship where the output gap depends on the expected output gap and the real interest rate, $i_t - E_t \pi_{t+1}$. Equation (3.2) is derived under assumption of monopolistic competition, with individual firms adjusting prices in a staggered, overlapping fashion. Also it is a forward-looking AS relationship where inflation depends on expected inflation and output gap. The relation of inflation and output gap in equation (3.2) is represented by general functional form, $F(\tilde{y}_t)$ since it is able to introduce a nonlinear relation. The nonlinearity in AS curve is specified by equation (3.3). This functional form has been used previously by Schaling (2004), Dolado et al. (2005), and Surcio (2003b), and Dolado, et al (2004), etc., because it can be a linear or a nonlinear AS curve depending on the value of the parameter (ψ) represented the degree of nonlinearity.[3] If the parameter is zero (i.e., $\psi = 0$), then the AS curve turns into the standard linear relation. But if the parameter is positive (i.e., $\psi > 0$), the relation is a nonlinear one as it allows the slope of the aggregate supply curve to be steeper at a higher level of inflation and output gap. It can be justified by the presence of a capacity constraint or some menu costs, downward wage rigidity, etc.

3.2.2 Policy Objectives of Central Banks

Following the literature, it is assumed that the central bank is trying to minimize the expected value of a loss function that depends on inflation, output, and interest rate.

$$Min_{\{i_t\}} \; E_{t-1} \sum_{j=0}^{\infty} \delta^j V_{t+j}, \tag{3.4}$$

where $0 < \delta < 1$ is the discount factor of central bank and V_t is the period loss function. The loss function is assumed to take the following linear-exponential (i.e., linex) form for inflation and output gap:

$$V_t = \left[\frac{\exp\{\varphi(\pi_t - \pi^*)\} - \varphi(\pi_t - \pi^*) - 1}{\varphi^2} \right] + \theta \left[\frac{\exp(\phi \tilde{y}_t) - \phi \tilde{y}_t - 1}{\phi^2} \right] + \frac{A}{2}(i_t - i_{t-1} - g)^2, \tag{3.5}$$

---

[3] It can be possible for the parameter to be negative (i.e., $\psi < 0$). This means that the relation between inflation and output gap is concave, but this case seems to be against the reality.

where π^* is target inflation rate and g is average (or target) of interest rate change.[4] The parameters θ and A are strictly positive and govern the relative weight that central bank places on output and interest rate stabilization relative to inflation stabilization, respectively. The parameter φ represents the degree of asymmetry with respect to inflation, while the parameter ϕ stands for that with respect to output gap.

The loss function in here is different with Surcio (2003b) in that we use $(i_t - i_{t-1})^2$ instead of $(i_t - i^*)^2$ where i^* is target rate.[5] This type of objective is used in Svensson (2000). This change has some interesting theoretical and empirical aspects. First, this loss function reflects the fact that the central bank control interest rate toward target rate almost perfectly as shown in Figure 3.1. Therefore, the loss from deviation from previous interest rate might be more important than the loss from deviation from the target rate in central bank's standpoint of view. Second, Surcio (2003b) transformed the reaction function derived from the FOC ad hoc when the partial adjustment behavior of interest rate introduced. However, in here we formally incorporate the adjustment cost into the period loss function. Last, the nominal interest rates show the highly persistent property in this sample period.[6] The possibility of spurious regression can be avoided since the reaction function derived from this objective is made of the change of interest rate as shown in Section 3.2.3.

[4] In other papers such as Svensson (2000) g is implicitly assumed to zero. In here it is introduced because first, it is more general and second, it make the monetary policy reaction function have a constant term. But as shown in empirical results constant term is not significant statistically.
[5] For the robustness, the monetary policy reaction function used in Surcio (2003b) derived and its estimating results are reported in Appendix.
[6] For the sample period-1998:9-2005:12 the ADF test can not reject the hypothesis of unit roots in all three different specifications, but this results depend on the sample period.

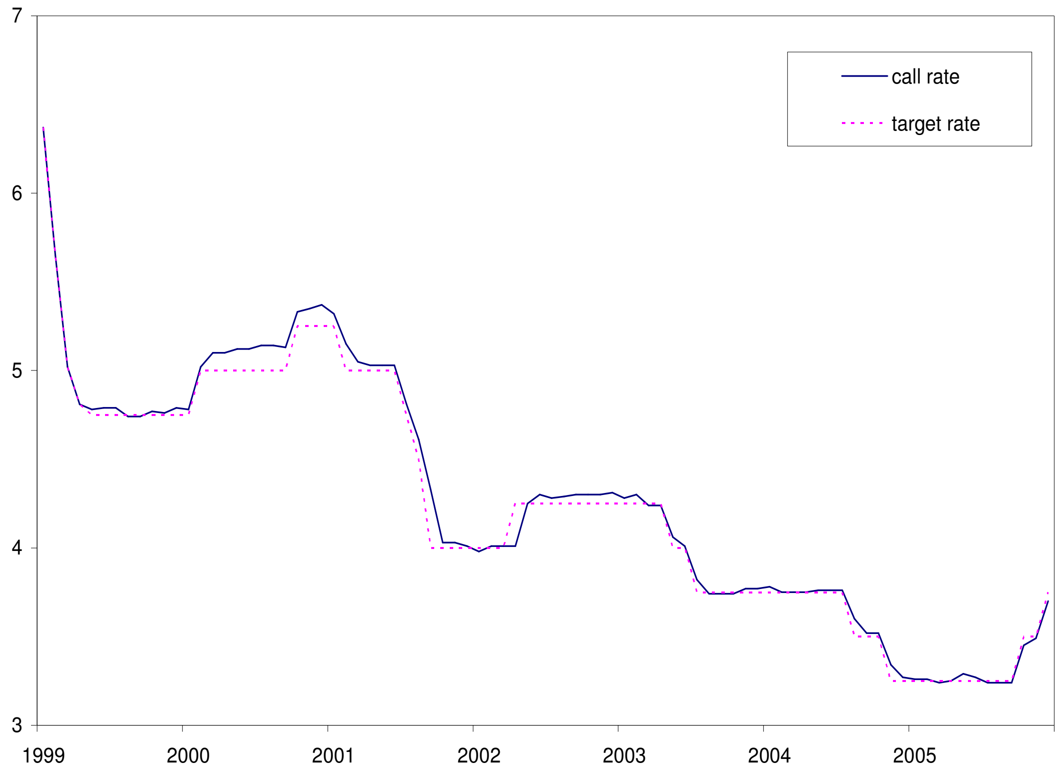


Figure 3.1 Observed and Target Call Rate

The linex functional form in equation (3.5) is frequently used in recent asymmetric optimal monetary policy literature, including Nobay and Peel (2003), Surcio (2003a, 2003b), Ruge-Murcia (2003) and Karagedikli and Lees (2004) etc., since it gives direction on the type of asymmetric preferences that serve as the metric to evaluate alternative monetary policies. Also the linex functional form nests the quadratic form as a special case, i.e., using L'Hopital's rule and differentiating twice with respect to

$\varphi = \phi = 0$, the conventional quadratic form, $V_t = \frac{1}{2}(\pi_t - \pi^*)^2 + \frac{\theta}{2}\tilde{y}_t^2 + \frac{A}{2}(i_t - i_{t-1} - g)^2$ could be recovered.[7] If φ (ϕ) is not zero, then the linex function weighs differently with respect to positive and negative deviations of inflation from target (output gaps). For example, if $\varphi > 0$, positive deviations of inflation from target are more costly than negative ones since the exponential term dominates the linear term in equation (3.5), while when $\varphi < 0$, negative deviations of inflation from target are more costly than positive ones because the linear term dominates the exponential term. As a same token, if $\phi < 0$, negative output gaps are more costly than positive ones, while positive output gaps are more costly than negative one when $\phi > 0$.

3.2.3 A Monetary Policy Reaction Function

The time frame of monetary policy making is that the interest rate has to be chosen before the realization of economic shocks is known with certainty to policy maker. This means that the innovations e_t^d and e_t^s are unknown at the time policy maker choose the nominal interest rate i_t. Also under discretion the policy maker takes expectations of future variables as given. Therefore, the problem of the policy maker is to choose the current interest rates and the sequence of future interest rates such as to minimize following function subject to the behavior of the economy:

[7] For example, $\lim_{\phi \to 0} \frac{\exp(\phi\tilde{y}_t) - \phi\tilde{y}_t - 1}{\phi^2} = \lim_{\phi \to 0} \frac{\tilde{y}_t \exp(\phi\tilde{y}_t) - \tilde{y}_t}{2\phi} = \lim_{\phi \to 0} \frac{\tilde{y}_t^2 \exp(\phi\tilde{y}_t)}{2} = \frac{\tilde{y}_t^2}{2}$.

$$E_{t-1}\left[\frac{\exp\{\varphi(\pi_t-\pi^*)\}-\varphi(\pi_t-\pi^*)-1}{\varphi^2}\right]+\theta E_{t-1}\left[\frac{\exp(\phi\tilde{y}_t)-\phi\tilde{y}_t-1}{\phi^2}\right]$$

$$+\frac{A}{2}(i_t-i_{t-1}-g)^2+\frac{\delta A}{2}E_{t-1}(i_{t+1}-i_t-g)^2+S_t \tag{3.6}$$

subject to

$$\tilde{y}_t=-\alpha i_t+f_t \ , \tag{3.7}$$

$$\pi_t=\beta\tilde{y}_t/(1-\beta\psi\tilde{y}_t)+g_t, \tag{3.8}$$

where $S_t \equiv E_{t-1}\sum_{j=2}^{\infty}\delta^j V_{t+j}$, $f_t \equiv E_t\tilde{y}_{t+1}+\alpha E_t\pi_{t+1}+e_t^d$ and $h_t \equiv \omega E_t\pi_{t+1}+e_t^s$ stand for the components of the model the monetary policy makers cannot control since they are not able to manipulate expectations directly.

The first order condition (FOC) for minimizing the loss of central bank is given by

$$-E_{t-1}\left[\frac{\exp\{\varphi(\pi_t-\pi^*)\}-1}{\varphi}\right]\frac{\alpha\beta}{(1-\beta\psi\tilde{y}_t)^2}-E_{t-1}\left[(\frac{\exp(\phi y_t)-1}{\phi})\right]\theta\alpha$$

$$+A(i_t-i_{t-1}-g)-\delta AE_{t-1}(i_{t+1}-i_t-g)=0. \tag{3.9}$$

Based on above equation, we can derive the various FOC specifications by combining different nonlinearity conditions. For example, when the condition $\varphi=\phi=\psi=0$ is imposed, the FOC (3.9) reduces to the following linear condition:

$$-\alpha\beta E_{t-1}(\pi_t-\pi^*)-\theta\alpha E_{t-1}\tilde{y}_t+A(i_t-i_{t-1}-g)-\delta AE_{t-1}(i_{t+1}-i_t-g)=0. \tag{3.10}$$

Also this equation can be derived directly from quadratic objective function of central bank and a linear system of the economy.

Now we proceed to an empirical specification. Equation (3.9) cannot be estimated directly since the equation is not linear in the parameters. Therefore, by taking the first order Taylor series expansion at $\varphi = \phi = \psi = 0$, the exponential terms and inverse function in equation (3.9) is approximated,[8] and then it can be written as

$$-\alpha\beta E_{t-1}(\pi_t - \pi^*) - \theta\alpha E_{t-1}\tilde{y}_t - \frac{\varphi\alpha\beta}{2}E_{t-1}(\pi_t - \pi^*)^2 - \frac{\phi\theta\alpha}{2}E_{t-1}\tilde{y}_t^2$$

$$-2\beta^2\psi\alpha E_{t-1}[(\pi_t - \pi^*)\tilde{y}_t] + A(i_t - i_{t-1} - g) - \delta A E_{t-1}(i_{t+1} - i_t - g) + \varepsilon_t = 0, \quad (3.11)$$

where ε_t represents the higher order of the Taylor series expansion. Solving for Δi_t and the expected inflation, output gap and interest rate change are replaced by actual values, then the nonlinear monetary policy reaction function can be written as follows:[9]

$$\Delta i_t = \delta\Delta i_{t+1} + c_0 + c_1\tilde{\pi}_t + c_2\tilde{y}_t + c_3\tilde{\pi}_t^2 + c_4\tilde{y}_t^2 + c_5(\tilde{\pi}_t\tilde{y}_t) + u_t, \quad (3.12)$$

which is linear in the coefficients, and for simplicity $\pi_t - \pi^*$ is defined as $\tilde{\pi}_t$ which stands for deviation of inflation from target or inflation gap. Also we get the following coefficients and error term condition related with equation (3.11) and (3.12):

$$c_0 \equiv (1-\delta)g, \quad c_1 \equiv \frac{\alpha\beta}{A}, \quad c_2 \equiv \frac{\theta\alpha}{A}, \quad c_3 \equiv \frac{\varphi\alpha\beta}{2A}, \quad c_4 \equiv \frac{\phi\theta\alpha}{2A}, \quad c_5 \equiv \frac{2\beta^2\psi\alpha}{A}, \text{ and}$$

---

[8] The Taylor series expansions of exponential function and inverse function follow:
$\exp(x) = 1 + x + \frac{x^2}{2!} + \frac{x^3}{3!} + \cdots$, $\frac{1}{(1-x)^r} = \sum_{k=0}^{\infty}\binom{r+k-1}{k}x^k \equiv \sum_{k=0}^{\infty}\binom{r+k-1}{r-1}x^k$, respectively.

[9] If we recursively solve equation (12) forward to yield

$$\Delta i_t = \sum_{i=0}^{\infty}\delta^i[c_0 + c_1\tilde{\pi}_{t+i} + c_2\tilde{y}_{t+i} + c_3\tilde{\pi}_{t+i}^2 + c_4\tilde{y}_{t+i}^2 + c_5(\tilde{\pi}_{t+i}\tilde{y}_{t+i}) + u_{t+i}].$$

The change of nominal interest rate depends on the discount value of future inflation gaps and output gaps and forecasting errors.

$$u_t \equiv -\frac{1}{A}\left\{\begin{array}{c}(\delta/A)(\Delta i_{t+1} - E_{t-1}\Delta i_{t+1}) + c_1(\tilde{\pi}_t - E_{t-1}\tilde{\pi}_t) + c_2(\tilde{y}_t - E_{t-1}\tilde{y}_t) \\ + c_3(\tilde{\pi}_t^2 - E_{t-1}\tilde{\pi}_t^2) + c_4(\tilde{y}_t^2 - E_{t-1}\tilde{y}_t^2) + c_5\{\tilde{\pi}_t\tilde{y}_t - E_{t-1}(\tilde{\pi}_t\tilde{y}_t)\} + \varepsilon_t\end{array}\right\}.$$

The error term is a linear combination of forecast errors and thus orthogonal to any variable in the information set available at t-1. Therefore, based on this orthogonality condition, the parameters in equation (3.12) can be estimated by the generalized method of moments (GMM). From the reduced coefficients, we can recover the asymmetric preferences over inflation gap and output gap such as $\varphi = 2c_3/c_1$, and $\phi = 2c_4/c_2$. However, two parameters of the weight on output gap relative to inflation (θ) and the degree of nonlinear in AS curve (ψ) are not recovered without knowing the parameter β in AS curve. The focus of this paper is to estimate the asymmetric preference parameters and tests whether those are statistically different with zero. Therefore, in here we do not try to recover ψ and just use the cross-product term as a control variable. In empirical aspect, we try to estimate the preference parameters in various specifications using the combination of nonlinearity conditions. For example, we can estimate equation (3.12) assuming either a linear or nonlinear AS curve.

3.3 Empirical Results

3.3.1 Data

In the empirical analysis, we use the monthly data for overnight call rate, yield of 5-year government bond, consumer price index (CPI), core consumer price index (core

CPI) and industrial production index (IPI), which are obtained from the Bank of Korea's Economic Statistics System.[10] The 12 month inflation rate is calculated based on CPI or core CPI. That is, $\pi_t = (\log P_t - \log P_{t-12}) \times 100$, where P_t is CPI or core CPI. Beginning the inflation targeting, the target was formed in terms of CPI. However, from year 2000, this was changed in terms of the core CPI which excluded certain non-grain agricultural products[11] and petroleum products from CPI. Therefore, we call the inflation based on CPI as CPI inflation, and the inflation based on core CPI since year 2000 as core inflation. As show in Figure 3.2, two inflation series moved similarly. But core inflation is more stable than CPI inflation. The target inflation is obtained from the published figures of the Bank of Korea, [12] and using it CPI and core inflation gaps are calculated as shown in Figure 3.3. Output gap series are constructed from the seasonally adjusted IPI using a quadratic trend method over 1998:1-2005:12. More specifically, output gap is measured as the actual output's percentage deviation from the trend output: $\tilde{y}_t = \{(q - q^*)/q^*\} \times 100$, where q and q^* are respectively the actual output and the trend output which are shown in Figure 3.4. Also the estimated output gap series are shown in Figure 3.5.

The sample is from 1998:4 to 2005:12. Since first five observations are used for central bank's information set, the actual sample used for estimating the reaction function of the central bank of Korea is 1998:9-2005:12. The descriptive statistics of the variables considered here are summarized in Table 3.2.

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[10] http://ecos.bok.or.kr/.

[11] Those consist of vegetables, fruits and other agricultural products.

[12] The inflation targets used in here are the middle figure of the range since the Bank of Korea publishes the target as a target range such as 2-4%.

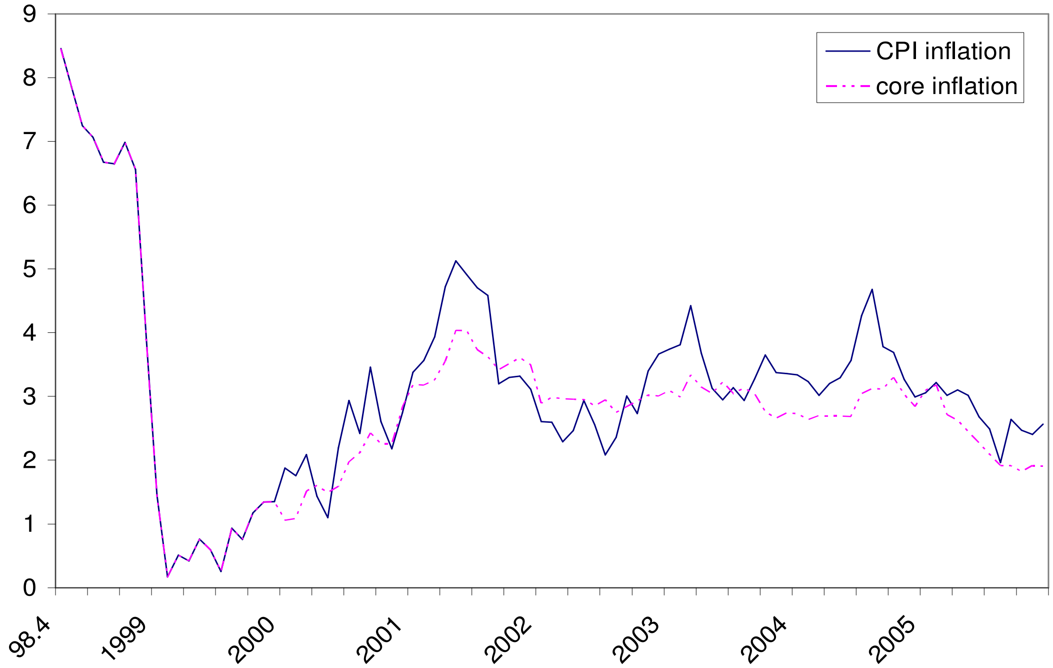

Figure 3.2 Annual CPI and Core Inflation

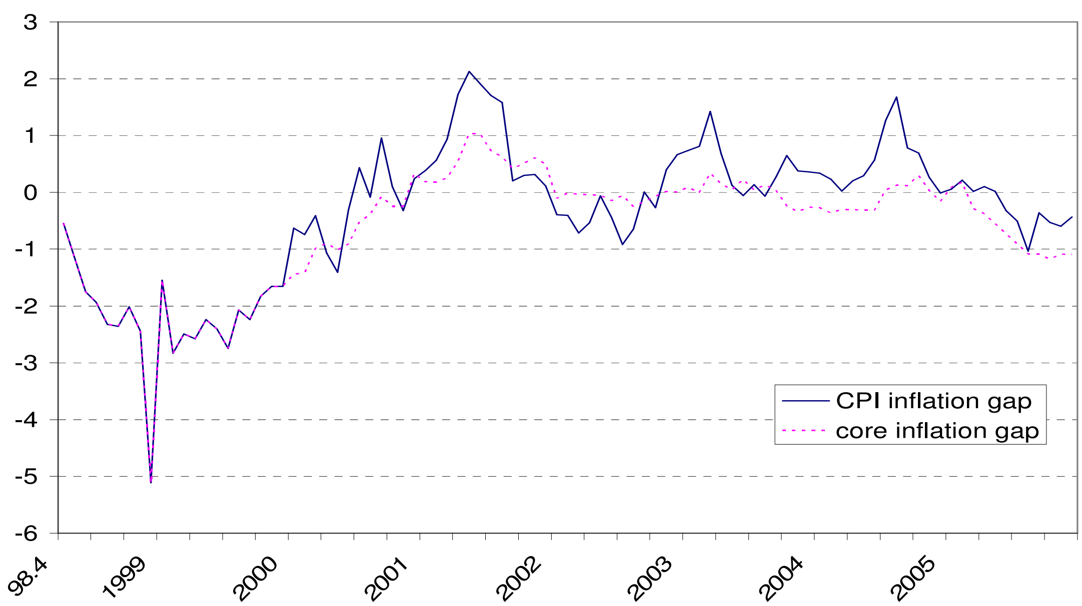

Figure 3.3 Annual CPI and Core Inflation Gaps

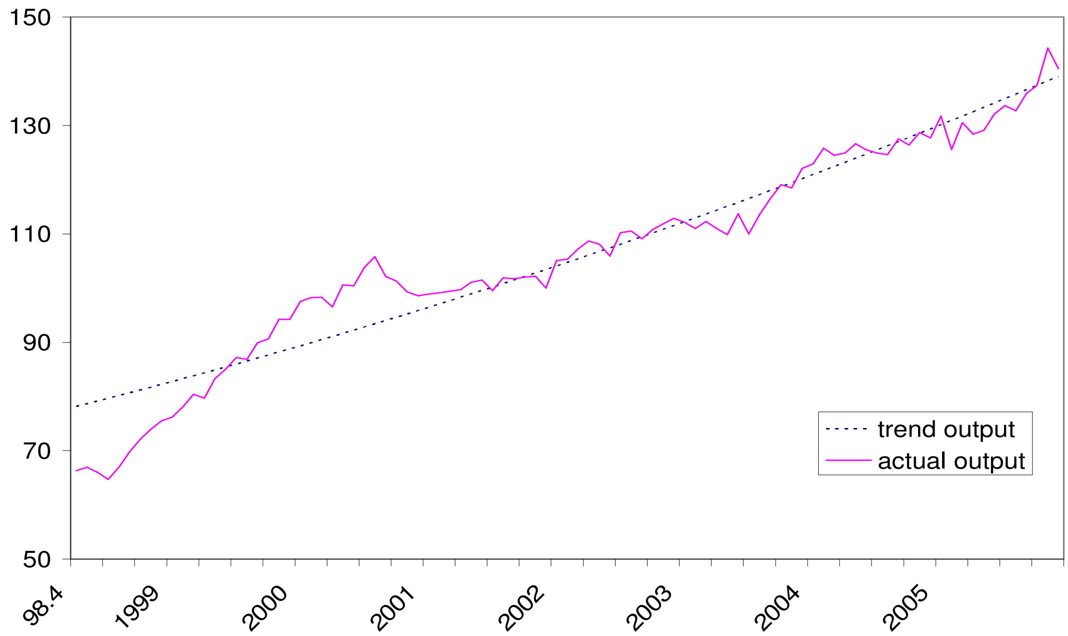

Figure 3.4 Actual Output and Estimated Output Trend Series

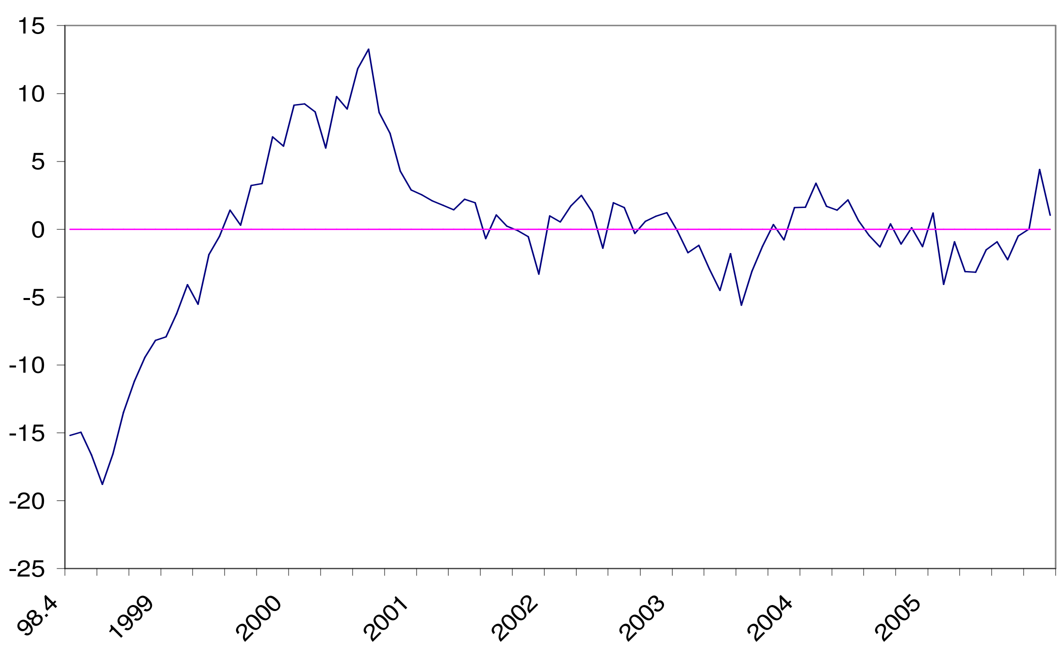

Figure 3.5 Estimated Output Gap Series

Table 3.2
Descriptive Statistics

	Call rate	Output gap	Inflation		
			CPI	Core	Target
Mean	4.430	0.558	2.926	2.641	3.205
Standard deviation	0.980	4.618	1.305	1.194	1.283

Sample period is from 1998:9 to 2005:12. CPI inflation is annual inflation rate calculated on consumer price index, while core inflation is calculated based on CPI until 1999:12 but after that based on Core CPI. Output gap is estimated using a quadratic trend.

3.3.2 Estimation Results

We estimate the equation (3.12) using the GMM, while a constant and five lags of inflation, output gap, call rate and the long-short interest rate spread[13] are included as instruments. Table 3.3 reports the results of estimating several reaction functions that allow for only an asymmetric response to inflation gap or/and the output gap, and also with/without a nonlinear AS curve when core inflation is used for central bank's target inflation. Over various specifications the estimates of the discounting factor of the Bank of Korea are all significant and are ranged from 0.835 to 0.923[14], which are relatively small number compared to the subjective rate of time discount in utility function of representative household generally used in standard real-business-cycle literature.[15]

[13] The long-short interest spread is constructed from yields of 5-year government bond and overnight call rate.
[14] If those values are converted into annual ones, then they are corresponding to 0.115 and 0.382, respectively.
[15] Walsh (2003) used the subjective rate of time discount in utility function as 0.989 for quarterly data (p. 75). This value is corresponding to 0.9567 for annual data. For monthly data the discount factor 0.996 is close to the corresponding annual discount rate.

Table 3.3
Reduced Form Estimates of Reaction Function

Model	δ	c_0	c_1	c_2	c_3	c_4	c_5	J	$\overline{R}^2$
Panel A: Linear reaction function									
(1)	0.855**	0.006	0.023**	0.005**	-	-	-	0.138	0.498
	(0.026)	(0.005)	(0.005)	(0.002)					
Panel B: Nonlinear reaction function with inflation gap only									
(2)	0.849**	0.010+	0.049**	0.006**	0.011	-	-	0.136	0.523
	(0.031)	(0.005)	(0.015)	(0.002)	(0.007)				
(3)	0.835**	0.009+	0.042**	0.005	0.008	-	-0.001	0.136	0.513
	(0.047)	(0.005)	(0.016)	(0.003)	(0.008)		(0.003)		
Panel C: Nonlinear reaction function with output gap only									
(4)	0.923**	0.005	0.024**	0.003+	-	0.0003	-	0.135	0.454
	(0.047)	(0.006)	(0.007)	(0.002)		(0.0003)			
(5)	0.865**	0.003	0.024**	-0.006	-	0.0005	-0.007+	0.120	0.415
	(0.042)	(0.007)	(0.009)	(0.006)		(0.0004)	(0.004)		
Panel D: Nonlinear reaction function with inflation gap and output gap									
(6)	0.923**	0.010+	0.050**	0.004+	0.012+	0.0003	-	0.133	0.482
	(0.059)	(0.006)	(0.015)	(0.002)	(0.007)	(0.0003)			
(7)	0.882**	0.007	0.045*	-0.006	0.009	0.0006	-0.007	0.122	0.442
	(0.055)	(0.007)	(0.021)	(0.007)	(0.009)	(0.0004)	(0.005)		

Estimation model is as follows: $\Delta i_t = \delta \Delta i_{t+1} + c_0 + c_1 \tilde{\pi}_t + c_2 \tilde{y}_t + c_3 \tilde{\pi}_t^2 + c_4 \tilde{y}_t^2 + c_5 (\tilde{\pi}_t \tilde{y}_t) + u_t$. The estimation period is from 1998:9 to 2005:12. Output gap is obtained from percentage change of actual and detrending industrial production index with quadratic trend equation for the sample period 1998:1-2005:12. The instrument sets include constant and five lags of core inflation, output gap, long-short interest rate spread and call rate. The superscript **, * and + denote the rejection of the null hypothesis that the true coefficient is zero at the 1 percent, 5 percent and 10 percent significance levels, respectively.

Another interesting fact is that when the cross-product term of inflation gap and output gap is specified in the reaction function, its coefficient has wrong sign, but marginally significant under asymmetry assumption over only output gap. Also the estimated coefficients of output gap become insignificant and they have wrong sign in 2 times out of 3 models.

Panel A in Table 3.3 reports the estimation results of the baseline linear reaction function which can be obtained by imposing the condition $c_4 = c_5 = c_6 = 0$ on equation (3.12). The coefficients of inflation and output gap are significant at 1 percent significance level. However, it appears that central bank of Korea was more aggressive toward inflation than output since the coefficient of inflation gap is 4 times bigger than that of output gap. Next, we consider the Bank of Korea's asymmetric preferences over inflation gap only with/without assuming a linear AS curve (i.e., $c_4 = 0$, or $\phi = 0$). Panel B depicts estimation results for a nonlinear response to deviations of inflation from target. First, if we consider a linear AS curve, the square of deviations of inflation from target is positive but not marginally significant at 10% significance level. Second, when the nonlinear AS curve is considered in the model, the coefficients of output gap become insignificant and also the fitness of the model is worse in term of adjusted R^2 compared to the model with a linear AS curve. In Panel C the estimation results considered only an asymmetry over output gap are reported. The square of output gap is positive but not significant regardless of the AS curve, so it indicates that the preference of the Bank of Korea is not asymmetric over output. Finally, when the possibility of asymmetric preferences over both inflation gap and output gap is considered which is more general specification (Panel D), the squares of inflation gap and output gap are not significant with the cross term of inflation gap and output gap. But when the cross term of inflation gap and output gap is dropped, then only the square of inflation gap is significant at 10% significance level. Those results seem to be for the hypothesis that the central bank of Korea may response asymmetrically to inflation gap but not to output gap.

However, before jumping to the conclusion we need to test statistically whether the asymmetry parameters are zero. In this end, we estimate the following equation for obtaining the asymmetric parameters and their standard deviations, and so we can test their significance statistically[16]:

$$\Delta i_t = \delta \Delta i_{t+1} + c_0 + \varphi\{(2a/\varphi)\tilde{\pi}_t + a\tilde{\pi}_t^2\} + \phi\{(2b/\phi)\tilde{y}_t + \phi\tilde{y}_t^2\} + c_6(\tilde{\pi}_t \tilde{y}_t) + u_t. \qquad (3.13)$$

Note that this equation is the same with equation (3.12) mathematically. This means that the estimation results should be same with Table 3.3, which is the result of estimating equation (3.12). Table 3.4 reports the estimated parameters of central bank's asymmetric preferences under various specifications. The asymmetry parameter over inflation gap is significant at 1% significance level when the AS curve is linear. The estimated parameter when considered both inflation gap and output gap is 0.483, which is very close with that when considered only inflation gap. Thus the Bank of Korea appears to possess asymmetric preferences with regard to the inflation gap, and since its sign is positive, we can infer that the Bank of Korea is out-weighting positive deviations of inflation from its target than negative ones in its loss function. However, the estimated

[16] Although the parameters of asymmetric preferences can be recovered from estimated coefficients of the reaction function using the relationship as $\varphi = 2c_3 / c_1$ and $\phi = 2c_4 / c_2$, the equation (13) should be estimated in order to test the significance of asymmetry parameters statistically. Different forms can also be used for this end.

Table 3.4
Estimates of the Asymmetric Policy Preferences

Model	Inflation gap (φ)	Output gap (γ)	J	$\overline{R}^2$
Panel A: Inflation gap only				
(2)	0.464**	-	0.136	0.522
	(0.161)			
(3)	0.385	-	0.136	0. 513
	(0.231)			
Penal B: Output gap only				
(4)	-	0.176	0.135	0.454
		(0.210)		
(5)	-	-0.174	0.120	0.415
		(0.112)		
Panel C: Inflation gap and output gap				
(6)	0.483**	0.147	0.133	0.482
	(0.157)	(0.212)		
(7)	0.411+	-0.198	0.122	0.442
	(0.238)	(0.145)		

Estimation model is as follows: $\Delta i_t = \delta \Delta i_{t+1} + c_0 + \varphi\{(2c_3/\varphi)\tilde{\pi}_t + c_3\tilde{\pi}_t^2\} + \phi\{(2c_4/\phi)\tilde{y}_t + c_4\tilde{y}_t^2\} + c_5(\tilde{\pi}_t\tilde{y}_t) + u_t$. The estimation period is from 1998:9 to 2005:12. Output gap is obtained from percentage change of actual and detrending industrial production index with quadratic trend equation for the sample period 1998:1-2005:12. The instrument sets include constant and five lags of core inflation, output gap, long-short interest rate spread and call rate. The superscript **, * and + denote the rejection of the null hypothesis that the true coefficient is zero at the 1 percent, 5 percent and 10 percent significance levels, respectively.

asymmetric parameters over output gap have relatively small values and their sign depends on existence of the cross term. However, the preference parameters are not significant at conventional significance level in all specification. Therefore, estimating results of asymmetric preference with regard to output gap seems to be against the asymmetric hypothesis.

In sum, this empirical evidence supports that while the preference over output of central bank of Korea is symmetry, but that over inflation is not symmetry. Specifically the bank responds aggressively positive inflation gaps compared to negative inflation gaps. This positive asymmetric preference over inflation can induce deflation bias, which is not big because the size of parameter is relatively small. Figure 3.6 graphs the estimated asymmetric preference over inflation gap compared to the standard quadratic representation assuming $\varphi = 0.483$. Note that since asymmetric parameter estimate over inflation gap is relatively small, the loss from inflation gaps is underweighted compared to that of the standard quadratic specification regardless of its sign.

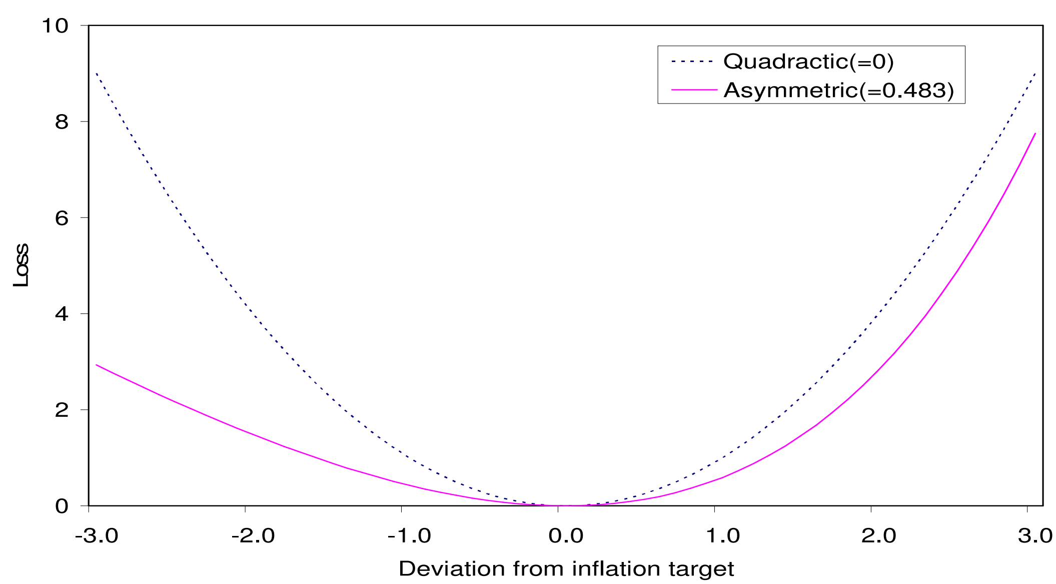

Figure 3.6 Asymmetric Preferences with Respect to Inflation of the Bank of Korea

3.3.3 Robustness Tests

The estimated discount factor in equation (3.12) is relatively small compared to the value normally assumed for representative household as discussed in Section 3.3.2, although the discount factor of central bank necessarily need not to be same with that of household. For the robustness, in here we re-estimate the reaction function by imposing some specific number of the discount factor on equation (3.12). Table 3.5 shows the estimation results when the restriction $\delta = 0.996$ is imposed. It is not surprising that the fitness of the model is deteriorated compared to the results in Table 3.3. One interesting thing is that the square of output gap becomes significant statistically although the estimated coefficients are very small. While the square of inflation gap also becomes significant statistically, some coefficients of inflation gap also become significant. When asymmetric preferences over inflation and output considered, the cross term of inflation gap and output gap is not significant statistically. However, the square terms of inflation gap and output gap are significant at the standard significance level in model 6. The estimation results of asymmetry parameters assuming $\delta = 0.996$ are reported in Table 3.6. The estimated asymmetric preference parameter over only inflation gap is statically significant regardless of function form of an AS curve as shown in Panel A. However, as reported in Panel B, if asymmetric preference over only output gap considered, the estimated preference parameter over output gap is not significant statistically. Finally, when asymmetric preferences over both inflation gap and output gap are specified in the

Table 3.5
Reduced Form Estimates of Reaction Function - $\delta = 0.996$

Model	δ	c_0	c_1	c_2	c_3	c_4	c_5	J	$\overline{R}^2$
Panel A: Linear reaction function									
(1)	0.996	0.012+	0.022**	0.003+	-	-	-	0.150	0.454
		(0.006)	(0.006)	(0.002)					
Panel B: Nonlinear reaction function with inflation gap only									
(2)	0.996	0.022**	0.091**	0.005**	0.030**	-	-	0.154	0.504
		(0.007)	(0.010)	(0.002)	(0.004)				
(3)	0.996	0.012*	0.041**	0.010**	0.012*	-	0.006**	0.141	0.458
		(0.006)	(0.011)	(0.003)	(0.005)		(0.002)		
Panel C: Nonlinear reaction function with output gap only									
(4)	0.996	0.004	0.020**	0.001	-	0.0005**	-	0.132	0.421
		(0.006)	(0.007)	(0.001)		(0.0002)			
(5)	0.996	-0.0002	0.021*	-0.007	-	0.0009*	-0.006	0.121	0.365
		(0.007)	(0.008)	(0.007)		(0.0004)	(0.005)		
Panel D: Nonlinear reaction function with inflation gap and output gap									
(6)	0.996	0.010+	0.053**	0.002	0.014*	0.0005**	-	0.131	0.452
		(0.006)	(0.014)	(0.001)	(0.006)	(0.0002)			
(7)	0.996	0.010	0.063**	-0.008	0.018+	0.0009*	-0.006	0.122	0.420
		(0.007)	(0.020)	(0.008)	(0.009)	(0.0004)	(0.005)		

Estimation model is as follows: $\Delta i_t = \delta \Delta i_{t+1} + c_0 + c_1 \tilde{\pi}_t + c_2 \tilde{y}_t + c_3 \tilde{\pi}_t^2 + c_4 \tilde{y}_t^2 + c_5 (\tilde{\pi}_t \tilde{y}_t) + u_t$. The estimation period is from 1998:9 to 2005:12. Output gap is obtained from percentage change of actual and detrending industrial production index with quadratic trend equation for the sample period 1998:1-2005:12. The instrument sets include constant and five lags of core inflation, output gap, long-short interest rate spread and call rate. The superscript **, * and + denote the rejection of the null hypothesis that the true coefficient is zero at the 1 percent, 5 percent and 10 percent significance levels, respectively.

Table 3.6
Estimates of the Asymmetric Policy Preferences - $\delta = 0.996$

Model	Inflation gap (φ)	Output gap (γ)	J	$\overline{R}^2$
Panel A: Inflation gap only				
(2)	0.667**	-	0.154	0.504
	(0.040)			
(3)	0.579**	-	0.141	0. 458
	(0.125)			
Panel B: Output gap only				
(4)	-	1.035	0.132	0.421
		(1.295)		
(5)	-	-0.238	0.121	0.365
		(0.157)		
Panel C: Inflation gap and output gap				
(6)	0.538**	0.672	0.131	0.452
	(0.134)	(0.517)		
(7)	0.565**	-0.242	0.122	0.420
	(0.142)	(0.170)		

Estimation model is as follows: $\Delta i_t = \delta \Delta i_{t+1} + c_0 + \varphi\{(2c_3/\varphi)\tilde{\pi}_t + c_3\tilde{\pi}_t^2\} + \phi\{(2c_4/\phi)\tilde{y}_t + c_4\tilde{y}_t^2\} + c_5(\tilde{\pi}_t\tilde{y}_t) + u_t$. The estimation period is from 1998:9 to 2005:12. Output gap is obtained from percentage change of actual and detrending industrial production index with quadratic trend equation for the sample period 1998:1-2005:12. The instrument sets include constant and five lags of core inflation, output gap, long-short interest rate spread and call rate. The superscript **, * and + denote the rejection of the null hypothesis that the true coefficient is zero at the 1 percent, 5 percent and 10 percent significance levels, respectively.

model, the asymmetric parameter over inflation gap is significant and positive regardless of functional form of an AS curve. But asymmetric parameter over output gap is not significant. Generally this result is consistent with that discussed in Section 3.3.2.

Although the central bank of Korea changed its target inflation as core inflation from CPI inflation since 2000, they might still care about the CPI inflation. Therefore, whether the asymmetric preference of central bank over inflation gap is robust to the change of targeted inflation is tested. Table 3.7 reports the results of estimating the reaction function using CPI inflation instead of core inflation. The cross term of inflation gap and output gap is not significant in all 3 specifications. With a linear AS curve the square terms of inflation gap and output gap are significant statistically. This result supports that the preference over inflation gap is still nonlinear. However, the preference over output gap also might be nonlinear, which is different with the results when core inflation is used.[17]

3.4 Conclusion

The central bank of Korea has been adopted explicit inflation targets since 1998. Therefore, this provides the interesting environment for testing whether the preferences of the Bank of Korea are consistent with the quadratic preference assumption that is standard within the monetary policy literature. Using linear-exponential function instead

[17] In here the results of estimating the asymmetric preferences over inflation gap or/and output gap are not reported because some results of estimating equation (3.13) do not converged into those of estimating equation (3.12).

Table 3.7
Reduced Form Estimates of Reaction Function – CPI Inflation

Model	δ	c_0	c_1	c_2	c_3	c_4	c_5	J	$\overline{R}^2$
Penal A: Linear reaction function									
(1)	0.863**	-0.001	0.016**	0.005**	-	-	-	0.148	0.497
	(0.027)	(0.005)	(0.005)	(0.002)					
Panel B: Nonlinear reaction function with inflation gap only									
(2)	0.887**	-0.014*	0.052**	0.007**	0.025**	-	-	0.149	0.480
	(0.037)	(0.006)	(0.009)	(0.002)	(0.003)				
(3)	0.847**	-0.019*	0.059**	0.006*	0.030**	-	-0.002	0.150	0.484
	(0.049)	(0.008)	(0.012)	(0.003)	(0.005)		(0.002)		
Panel C: Nonlinear reaction function with output gap only									
(4)	0.517**	-0.002	0.011+	0.016**	-	-0.0014**	-	0.137	0.614
	(0.071)	(0.005)	(0.007)	(0.002)		(0.0003)			
(5)	0.871**	-0.006	0.013+	0.0003	-	0.0002	-0.003	0.146	0.470
	(0.041)	(0.006)	(0.007)	(0.0004)		(0.0004)	(0.003)		
Panel D: Nonlinear reaction function with inflation gap and output gap									
(6)	0.726**	-0.015*	0.054**	0.012**	0.026**	-0.0006*	-	0.154	0.524
	(0.075)	(0.007)	(0.011)	(0.003)	(0.003)	(0.0003)			
(7)	0.532**	0.005	0.005	0.009*	-0.009	-0.0009+	-0.004	0.119	0.532
	(0.076)	(0.008)	(0.011)	(0.005)	(0.005)	(0.0005)	(0.003)		

Estimation model is as follows: $\Delta i_t = \delta\Delta i_{t+1} + c_0 + c_1\tilde{\pi}_t + c_2\tilde{y}_t + c_3\tilde{\pi}_t^2 + c_4\tilde{y}_t^2 + c_5(\tilde{\pi}_t\tilde{y}_t) + u_t$. The estimation period is from 1998:9 to 2005:12. Output gap is obtained from percentage change of actual and detrending industrial production index with quadratic trend equation for the sample period 1998:1-2005:12. The instrument sets include constant and five lags of CPI inflation, output gap, long-short interest rate spread and call rate. The superscript **, * and + denote the rejection of the null hypothesis that the true coefficient is zero at the 1 percent, 5 percent and 10 percent significance levels, respectively.

of the standard quadratic function this paper examines the asymmetric preference of the Bank of Korea with regard to inflation gap and output gap under the new Keynesian economic framework. Under this framework, the nonlinear monetary policy reaction function is derived from the optimization behavior of central bank to minimize its loss. In addition, we introduce the possibility that AS curve is not linear, which also induces the nonlinear monetary reaction function.

We estimate the reaction function and then recover the asymmetric preference parameter over inflation gap and output gap of the Bank of Korea in sample 1998:9-2005:12. In summery, with caution we conclude that the AS curve may not be nonlinear form because the cross-product term of inflation gap and output gap is not significant and its sign is wrong. Second, with a linear AS curve the Bank of Korea have an asymmetric preference over inflation gap but not output gap. Third, since the sign of the preference parameter is positive, it appears that the BOK has the more weight on positive deviations of inflation from target than on negative ones, which supports the hypothesis of precautionary demand for inflation.

CHAPTER IV

ASYMMETRIC ADJUSTMENT IN AGGREGATE DIVIDENDS

4.1 Introduction

After Lintner (1956) proposed the well-known behavioral model of dividend policy, a number of papers try to model the behavior of dividend in both disaggregate and aggregate level. For aggregate level, Marsh and Merton (1987) developed the dynamic behavior model of aggregate dividend which has a feature that dividend is adjusted to its long run target dividend (i.e., error correction term). Also stock prices instead of accounting earnings are used to measure permanent earnings. Garrett and Priestley (2000) generalize Lintner model by introducing the manager's optimizing behavior. In their model, they assume that managers have the target dividends, and that adjustment costs lead firms not to adjust completely to the target dividends. Therefore, it postulates that there are costs associated with adjusting dividends and also costs associated with deviating from the target dividend. Under this linear quadratic objective framework, managers are trying to minimize these costs by setting the current dividends. They derived and estimated the generalized error correction model of dividend behavior which can embrace both Lintner model and Marsh and Merton model.

While most papers have analyzed the dividend behavior in the linear functional form, there is some possibility that adjustment of dividends may not be a linear process. One stylized fact of dividends behavior documented by Lintner (1956) is that most managers are willing to avoid making changes in dividends that stand a good chance of having to be reversed within the near future. Another important feature of dividends behavior is that there is asymmetry in dividend payments due to, for instance, a reluctance to cut dividends. For example, Yoon and Starks (1995) document the evidence that there is an asymmetry between dividend increases and dividend decreases at the individual firm level. Jalilvand and Harris (1984) examined the process of partial adjustment by allowing speeds of adjustment to vary by firm and over time depending on the size of firm and capital market conditions such as interest rates and stock prices. In addition, Marsh and Merton (1987) supported asymmetric adjustment of dividends. The deterministic component of their model is specified to be reflected "the standard text book proposition that, if the current payout is high relative to permanent earnings and therefore the retention rate is low, then dividends per share will be expected to grow more slowly than if the current payout were lower and the retention rate were corresponding higher."(p. 9).

As Garrett and Priestley (2000) pointed out, in Lintner's partial adjustment dividends model, there is no mechanism connecting the cost of adjustment with previous dividends and the cost of deviations from the target. [18] However, Managers' preferences about adjustment cost might be different depending on the degree or the sign of dividends deviation from the target. In other words the degree of the persistence of dividends change may be different depending on the size of dividends deviation from the target.

[18] They point out that Lintner model has an unattractive feature that adjustment of dividends is penalized irrespective of whether the adjustment brings the actual value closer to the target, and so in their extended model, movement toward the target lowers costs even if adjustment costs prevent a complete movement to the target.

In this chapter, we examine whether adjustment cost of dividends depends on the state of previous difference between dividends and the target. For example, adjustment cost of dividends faced by managers may be low (or high) when the previous dividends are above the target than when those are below the target. In terms of the speed of adjustment, when the previous dividends are above the target, the speed of adjustment may be faster (or slower) than when those are below the target. However, we are only accepting the hypothesis that the adjustment cost is regime dependent, but we are not going to constrain the managers' preferences and either to set the threshold point as zero. Rather, those are allowed to be determined in econometric analysis of data.

Under the regime dependent adjustment cost hypothesis, we are able to induce the model in which the changes of dividends follow the nonlinear error correction process from managers' optimization problem minimizing the costs of dividends adjustment. Therefore, the threshold vector error correction model (VECM) proposed by Hansen and Seo (2002) can be used to estimate the nonlinear adjustment process of dividends. Specifically, this paper analyzes the asymmetric adjustment behavior of the aggregate dividends in stock market with the threshold VECM, which allows for nonlinear adjustment cost and cointegration relationship. Dividends are corresponding of S&P 500 Stock Price Index. Real stock prices are used for a proxy for the target. We find significant evidence of threshold effect in dividends adjustment when stock prices are used a proxy for the target dividend over 1871q1-2004q2. This suggests that the adjustment costs are regime-dependent. We also find that when the difference of dividends and target is higher than the threshold, adjustment cost, which has the inverse relation with adjustment speed, is much smaller than that when it is lower.

The rest of this chapter is organized as follows. In Section 4.2, the nonlinear version of Lintner model is derived from a linear quadratic cost framework proposed by Garrett and Priestley (2000). Section 4.3 outlines the econometric methods to assess the asymmetric behavior of dividends. The empirical results are reported in Section 4.4 and the last section concludes.

4.2 The Model

In this paper, we follow Garrett and Priestley's (2000) framework, which assumes that there are two costs associated with adjusting dividends and with deviating from the target dividend.[19] The basic assumption of their model is that managers or dividend policy makers are trying to minimize the total cost (or loss) from the costs of adjusting dividends toward the target dividend and adjustment costs, which are quadratic because setting dividends above the target is as costly as setting ones below the target. The Lintner model in Garrett and Priestley framework can be derived from the following objective function:

$$\underset{\{d_t\}}{Min} \quad L = \phi(d_t - d_t^*)^2 + \varphi(\Delta d_t - g)^2, \tag{4.1}$$

[19] Basically Garrett and Priestley (2000) assume that adjustment costs lead firms not to adjust completely to the target dividends in every time period but instead to follow a pattern of partial adjustment. However, Jalilvand and Harris (1984) pointed out market imperfections as one source of partial adjustment.

where d_t is observed (log) dividends, d_t^* is target (log) dividends, $\Delta d_t = d_t - d_{t-1}$, and g is a normal growth rate. ϕ and φ are positive weighting parameters, which stand for deviation cost from the target and adjustment cost from the previous period, respectively.

If we solve equation (4.1) with respect to d_t, then the Lintner's partial adjustment model can be derived as follow:

$$\Delta d_t = (\frac{\varphi}{\phi+\varphi})g + (\frac{\phi}{\phi+\varphi})\{d_t^* - d_{t-1}\}. \tag{4.2}$$

Based on the possibility of asymmetric behavior of dividends documented by several literature, we assume that adjustment cost (i.e., φ) imposed by managers is regime dependent based on the degree or sign of the previous period's deviation from the target. Considering this asymmetric adjustment cost, the manager's problem we consider can be represented as follows:

$$\begin{aligned} \underset{\{d_t\}}{Min} \quad L = {} & \phi(d_t - d_t^*)^2 \\ & + \varphi_1(\Delta d_t - g)^2 \cdot 1(d_{t-1} - d_{t-1}^* \leq \gamma) \\ & + \varphi_2(\Delta d_t - g)^2 \cdot 1(d_{t-1} - d_{t-1}^* > \gamma), \end{aligned} \tag{4.3}$$

where $1(\cdot)$ is indicator function, and γ is threshold parameter.

Differencing equation (4.3) with respect to d_t and arranging it as d_t, then it yields

$$\begin{aligned} d_t = {} & \{(\frac{\varphi_1}{\phi+\varphi_1})g + (\frac{\phi}{\phi+\varphi_1})d_t^* + (\frac{\varphi_1}{\phi+\varphi_1})d_{t-1}\} \cdot 1(d_{t-1} - d_{t-1}^* \leq \gamma) \\ & + \{(\frac{\varphi_2}{\phi+\varphi_2})g + (\frac{\phi}{\phi+\varphi_2})d_t^* + (\frac{\varphi_2}{\phi+\varphi_2})d_{t-1}\} \cdot 1(d_{t-1} - d_{t-1}^* > \gamma). \end{aligned} \tag{4.4}$$

Next, d_{t-1} is subtracted from both sides in equation (4.4) and rearranging it gives

$$\Delta d_t = \{(\frac{\varphi_1}{\phi+\varphi_1})g + (\frac{\phi}{\phi+\varphi_1})d_t^* - [1-(\frac{\varphi_1}{\phi+\varphi_1})]d_{t-1}\} \cdot 1(d_{t-1} - d_{t-1}^* \le \gamma)$$

$$+\{(\frac{\varphi_2}{\phi+\varphi_2})g + (\frac{\phi}{\phi+\varphi_2})d_t^* - [1-(\frac{\varphi_2}{\phi+\varphi_2})]d_{t-1}\} \cdot 1(d_{t-1} - d_{t-1}^* > \gamma). \tag{4.5}$$

Also equation (4.5) can be written as

$$\Delta d_t = [(\frac{\varphi_1}{\phi+\varphi_1})g - (\frac{\phi}{\phi+\varphi_1})\{d_{t-1} - d_t^*\}] \cdot 1(d_{t-1} - d_{t-1}^* \le \gamma)$$

$$+[(\frac{\varphi_2}{\phi+\varphi_2})g - (\frac{\phi}{\phi+\varphi_2})\{d_{t-1} - d_t^*\}] \cdot 1(d_{t-1} - d_{t-1}^* > \gamma), \tag{4.6}$$

which is a nonlinear partial adjustment model where the coefficient of $d_{t-1} - d_t^*$ is $\phi/(\phi+\varphi_1)$ when $d_{t-1} - d_{t-1}^* \le \gamma$, but $\phi/(\phi+\varphi_2)$ when $d_{t-1} - d_{t-1}^* > \gamma$. However, If the adjustment cost does not depend on the state of actual dividend's deviation from the target in previous period (i.e., $\varphi = \varphi_1 = \varphi_2$), equation (4.6) will be reduced into a linear partial adjustment model (i.e., equation (4.2)).

Now in order to make equation (4.6) be a workable model, the target dividend generating process is specified. We assume that the target follows Martingale process, then it can be written such as

$$d_t^* = d_{t-1}^* + \varepsilon_t, \tag{4.7}$$

where $E_{t-1}(\varepsilon_t) = 0$.

Combined with equation (4.7), equation (4.6) can be expressed as

$$\Delta d_t = [(\frac{\varphi_1}{\phi+\varphi_1})g - (\frac{\phi}{\phi+\varphi_1})\{d_{t-1} - d_{t-1}^*\}] \cdot 1(d_{t-1} - d_{t-1}^* \le \gamma)$$

$$+[(\frac{\varphi_2}{\phi+\varphi_2})g-(\frac{\phi}{\phi+\varphi_2})\{d_{t-1}-d^*_{t-1}\}]\cdot 1(d_{t-1}-d^*_{t-1}>\gamma)$$

$$+[(\frac{\phi}{\phi+\varphi_1})\cdot 1(d_{t-1}-d^*_{t-1}\leq\gamma)+(\frac{\phi}{\phi+\varphi_2})\cdot 1(d_{t-1}-d^*_{t-1}>\gamma)]\cdot\varepsilon_t. \qquad (4.8)$$

Also it can be written as a concise form:

$$\Delta d_t=[\mu_1-\alpha_1(d_{t-1}-d^*_{t-1})]\cdot 1(d_{t-1}-d^*_{t-1}\leq\gamma)$$

$$+[\mu_2-\alpha_2(d_{t-1}-d^*_{t-1})]\cdot 1(d_{t-1}-d^*_{t-1}>\gamma)+u_t, \qquad (4.9)$$

where $E_{t-1}(u_t)=0$. Also we get the following coefficients and error term conditions from the relation between equation (4.8) and equation (4.9):

$$\mu_1=\frac{\varphi_1}{\phi+\varphi_1},\ \ \alpha_1=\frac{\phi}{\phi+\varphi_1},\ \ \mu_2=\frac{\varphi_2}{\phi+\varphi_2},\ \ \alpha_2=\frac{\phi}{\phi+\varphi_2},\ \text{and}$$

$$u_t=[(\frac{\phi}{\phi+\varphi_1})\cdot 1(d_{t-1}-d^*_{t-1}\leq\gamma)+(\frac{\phi}{\phi+\varphi_2})\cdot 1(d_{t-1}-d^*_{t-1}>\gamma)]\cdot\varepsilon_t.$$

Equation (4.9) represents the standard nonlinear error correction process, which adjustment coefficients depend on the state of $d_{t-1}-d^*_{t-1}$ i.e., whether $d_{t-1}-d^*_{t-1}\leq\gamma$ or $d_{t-1}-d^*_{t-1}>\gamma$.

4.3 Econometric Methods

This section develops econometric models that can be used to estimate the nonlinear behavior of dividend policy. One problem in estimating equation (4.9) is that it cannot be estimated in its current form because target dividends (i.e., d^*_{t-1}) are not observable.

Some literature uses the ratio of the fundamentals as a proxy for the target dividend. For example, Marsh and Merton (1987) defined the target as a linear function of the lagged log prices, while Garrett and Priestley (2000) specified that log target dividends are a linear function of log prices and log permanent earnings. In sum, the target can be calculated as constant payout ratio times the fundamentals. Following the literature, we also assumed that some ratio of the fundamentals (f_{t-1}) will be a good proxy for the log target (i.e., $d^{*}_{t-1} = \beta f_{t-1}$). The fundamentals we consider are log prices (p_{t-1}).[20]

We denote $x_t = (d_t, p_t)'$ and then the nonlinear (or threshold) vector error correction model (VECM) can be defined as follows:

$$\Delta x_t = [\mu_1 + \alpha_1 \omega_{t-1} + \sum_{i=1}^{k} \Gamma_{1,t-i} \Delta x_{t-i}] \cdot 1(\omega_{t-1} \le \gamma)$$

$$+ [\mu_2 + \alpha_2 \omega_{t-1} + \sum_{i=1}^{k} \Gamma_{2,t-i} \Delta x_{t-i}] \cdot 1(\omega_{t-1} > \gamma) + e_t, \qquad (4.10)$$

where $E_{t-1}(e_t) = 0$ and $1(\cdot)$ is indicator function. The long-run relationship is defined as $\omega_{t-1} = d_{t-1} - \beta p_{t-1}$, which is stationary as discussed by Engel and Granger (1987).

This chapter uses Hansen and Seo (2002)'s a grid-search algorithm for estimating the threshold VECM when the cointegrating vector is unknown. In here we briefly explain this algorithm. If we define the parameter vector $\theta = (\mu, \alpha, \Gamma_{t-1}, \cdots, \Gamma_{t-k})'$, and fix β and γ, then the threshold VECM can be estimated by linear regression

[20] Marsh and Merton (1987) distinguished economic earnings and accounting earnings. The former can be represented by stock prices under certain constraints.

$$\theta_1(\beta,\gamma) = [\sum_{i=1}^{k} z_t z_t^{'} 1(\omega_{t-1} \le \gamma)]^{-1} \sum_{i=1}^{k} z_t x_t^{'} 1(\omega_{t-1} \le \gamma),$$

$$\theta_2(\beta,\gamma) = [\sum_{i=1}^{k} z_t z_t^{'} 1(\omega_{t-1} > \gamma)]^{-1} \sum_{i=1}^{k} z_t x_t^{'} 1(\omega_{t-1} > \gamma),$$

$$\sum(\beta,\gamma) = \frac{1}{n}\sum_{i=1}^{k} e_t(\beta,\gamma) e_t^{'}(\beta,\gamma),$$

where $z_t^{'} = (1, \omega_{t-1}, \Delta x_{t-1}, \cdots, \Delta x_{t-k})^{'}$ and $e_t(\beta,\gamma)$ is in equation (4.10) with linear estimates for fixed β and γ.

1. Make a grid on [γ_L, γ_U] and [β_L, β_U] based on the linear estimate $\tilde{\beta}$.
2. For each value of (β,γ) on this grid, calculate $\theta_1(\beta,\gamma)$, $\theta_2(\beta,\gamma)$ and $\sum(\beta,\gamma)$.
3. Get $(\hat{\beta},\hat{\gamma})$ as the values of (β,γ) on this grid which yields the lowest value of $\log|\sum(\beta,\gamma)|$.
4. Find $\sum(\hat{\beta},\hat{\gamma})$, $\theta_1(\hat{\beta},\hat{\gamma})$, $\theta_2(\hat{\beta},\hat{\gamma})$ and $e_t(\hat{\beta},\hat{\gamma})$.

For testing, the threshold parameter cannot be identified under the null hypothesis, and as a result the standard methods cannot be applied. Therefore, we use the SupLM statistic defined in Hansen and Seo (2002), which does not depend on the nuisance parameter. [21]

$$SupLM_n = \underset{\gamma \in [\gamma_L, \gamma_u]}{Sup} LM_n,$$

where γ_L and γ_u satisfy $P(\omega_{t-1} \le \gamma_L) = p$ and $P(\omega_{t-1} > \gamma_u) = 1 - p$, respectively. The threshold parameter γ_L is the pth percentile of deviation from the target, and γ_u is the

[21] The details are referred to Hansen and Seo (2002).

(1-p)th percentile. Depending on the degree of freedom, p-value can be set at 0.05, 0.10, or 0.15. The SupLM statistic has a nonstandard asymptotic distribution. Therefore, the bootstrapping p-values are calculated, and if the bootstrapping p-values are smaller than the size chosen, we reject the null hypothesis.

4.4 Empirical Results

The data set used in the estimation consists of the Standard and Poor's 500 Composite Stock Price Index and corresponding dividends in quarterly frequency.[22] This frequency is reflected the fact that public companies usually pay dividends four times a year. The data is obtained from Robert Shiller's web site at http:// www.econ.yale.edu. The sample spans from 1871q1 to 2004q2. All series are log real values which are calculated by taking the log of the relevant variable divided by the consumer price index, and they are shown in Figure 4.1. One distinguishing feature is that the dividend series show more smooth movement than price series during entire period.

Since econometric model is based on the existence of cointegration relation between dividends and their target, we first analyze their stationary properties using the Augmented Dicky and Fuller (ADF) test. As shown in Table 4.1, the unit root hypothesis on stock prices can not be rejected both with drift and with drift and trend at 5% significance level, respectively, but the unit root hypothesis on dividends cannot be

[22] The quarterly data is extracted at quarterly frequency from Schiller's monthly data. The details are referred to Shiller (2000).

rejected when only drift is included. Although there is some controversy about the stationarity of these variables, in this paper dividends are assumed to follow the nonstationary process.

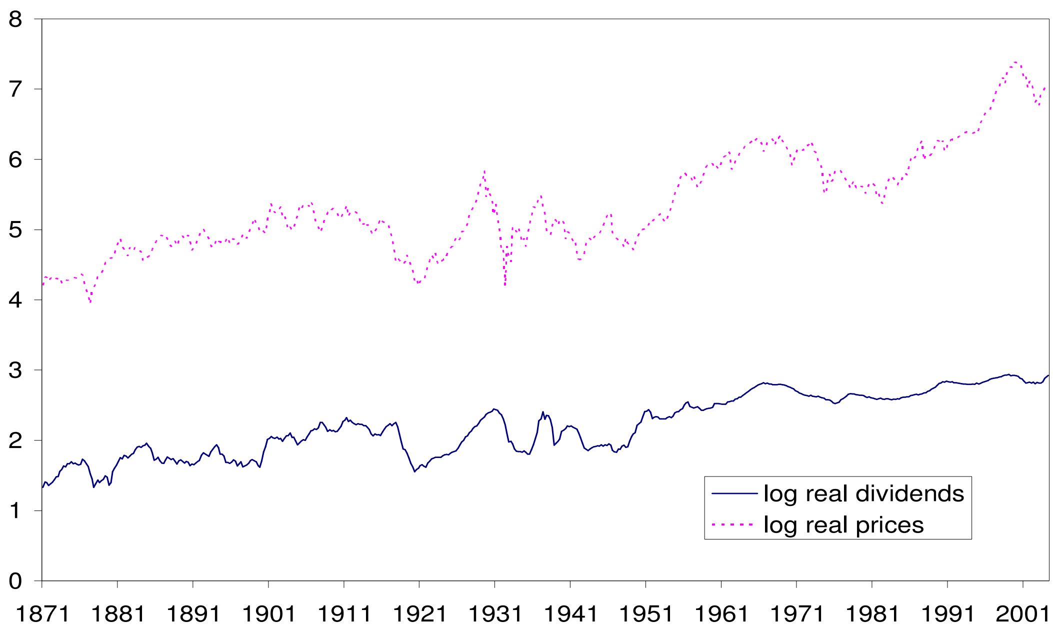

Figure 4.1. Real Dividends and Real Prices Series

Table 4.1
ADF Tests for Unit Root

Variable name	d_t	p_t
With drift	-1.628(1)	-0.907(0)
With drift and trend	-4.214(2)**	-2.386(0)

Dividends (d_t) and stock prices (p_t), and earnings(y_t) are all in logs. The lag order for the tests was determined by SIC in maximum lags with 18 for quarterly. Critical values are -3.442(1%), -2.867(5%) with drift and -3.975(1%), -3.418(5%) with drift and trend for quarterly. ** indicates rejection of null hypothesis at 1% significance level.

If stock prices are good proxies for the target, those variables should be cointegrated with dividends since dividends can not drift away from their target in the long-run. Table 4.2 shows the cointegration test results between dividends and stock prices by Johansen's (1988) reduced rank cointegration test. The likelihood ratio statistics reject the null hypothesis of no cointegration at 1% significance level regardless of target proxies.[23] Thus, we find a long-run relationship between those variables. This long-run relationship supports the assumption that stock prices might be good proxy variables for the target dividends.

Table 4.2
Cointegration Tests

Model	(d_t, p_t)	(d_t, y_t)
LR (H_0 :rank=0)	33.296**	58.772**
LR (H_0 :rank=1)	0.661	2.087
Lag length selected	2	2

The lag order for the tests was determined by SIC in maximum lags with 12. Critical values are 15.495 with H_0: rank=0 and 3.841 with H_0: rank=1. ** (*) indicates rejection of null hypothesis at 1% (5%) significance level.

Next, we estimate the linear vector error correction model without lag $(k = 0)$, which is consistent with the theoretical model derived in Section 4.2, for dividends and stock

[23] Since unit root hypothesis of dividends but prices is rejected with drift and trend, the cointegration test results might be come from the stationarity of dividends. Therefore we impose the restriction of (1, 0) on cointegrating vector, and then test this restriction statistically. Because Likelihood Ratio statistic is calculated at 29.653 with p-value 0.000, that restriction is rejected at 1% significance level. This supports that dividends and prices are cointegrated in the sample.

prices.[24] Table 4.3 reports the estimation results. The adjustment coefficient of dividends is significant at a standard significance level, while that of stock prices is not significant. This suggests that the deviation of dividends from their target in previous period is adjusted by the change of dividends but by the change of prices. No response of stock prices to the deviation seems to support the Miller and Modigliani (1961) hypothesis that the dividend policy does not affect the value of firm or stock prices.

Table 4.3
Estimation of Linear VECM

Model	VECM without lag		VECM with 1 lag	
Dependent variable	Δd_t	Δp_t	Δd_t	Δp_t
α	**-0.050**	0.007	**-0.047**	0.007
(s.e.)	(0.010)	(0.032)	(0.009)	(0.028)
μ	**-0.043**	0.011	**-0.038**	0.011
(s.e.)	(0.009)	(0.031)	(0.008)	(0.031)
Δd_{t-1}			**0.399**	0.112
(s.e.)			(0.072)	(0.108)
Δp_{t-1}			-0.016	-0.017
(s.e.)			(0.022)	(0.078)
β	0.585		0.572	
Likelihood	2,517.733		2,563.406	

The bold number indicates that it is significant statistically at 5% significant level.

[24] We also estimate the linear VECM model for dividends and earnings. However the estimation results do not report in here because the test of nonlinear adjustment hypothesis is rejected as shown in Table 4.4.

Some papers such as Tsai (2005) pointed the adjustment cost as main factor in the sluggish adjustment of dividends. We also can infer the relative importance of deviation cost and adjustment cost based on the relationship between adjustment coefficient in a linear VECM with the relation with parameters of theoretical model (i.e., $\alpha = \phi/(\phi + \varphi)$). Since the adjustment coefficient of dividends is only 0.05, we can induce a fact that the cost of adjustment (φ) from the previous dividends is very bigger than the cost of deviation (ϕ) from the target. This confirms the finding of Tsai (2005).[25]

Before estimating the nonlinear model of dividends adjustment, we need to test the hypothesis of a nonlinear adjustment of dividends. To allow for regime dependent adjustment cost of dividends, we use the threshold vector error correction model proposed by Hansen and Seo (2002).[26] The test results are reported in Table 4.4. The tests for threshold effects support the hypothesis of nonlinear adjustment in dividend determination process when stock prices are used for the target. For example, SupLM statistic for the tests of nonlinear adjustment in the model (d_t, p_t) is calculated at 17.236 with a bootstrapping *p*-value of 0.019. The tests are based on the threshold VECM without lag and the trimming parameter *p*=0.10. The bootstrapping p-values are calculated on the linear error correction model with 1000 bootstrapping replications.

[25] If two costs are equal, then adjustment coefficient of dividends will be 0.5. But the cost of adjustment is 19 times larger than the cost of deviation from target since it is 0.05. Tsai (2005) finds the evidence that adjustment cost is four times bigger than deviation cost in different model specification from monthly data.
[26] The gauss program used in this paper is generously provided by Byeongseon Seo.

Table 4.4
Tests for Nonlinear Adjustment

Model	Without lag		With one lag	
Boottrapping method	Fixed	boot	Fixed	boot
SupLM	17.236		21.729	
5% c.v.	15.246	14.388	20.769	21.236
p-value	0.021	0.019	0.035	0.042

Table 4.5 shows the estimation results of the threshold VECM, which is estimated by maximum likelihood estimation at the VAR lag-length 1 $(k=0)$. Standard errors are calculated from the heteroskedasticity-robust covariance estimator. The trimming parameter p is set at 0.10. The threshold estimate is -1.125 and $P(\omega_{t-1} \leq \gamma_L) = p$, and $P(\omega_{t-1} > \gamma_u) = 1-p$ are estimated at 0.818, and 0.182, respectively. In Figure 4.2 the upper and lower regime are depicted by time. The adjustment speed coefficients of the changes of dividends are -0.017 in regime 1 (lower regime) and -0.183 in regime 2 (upper regime), respectively. As expected in theoretical model the sign of the coefficients is negative in both regimes. However, only adjustment coefficient in regime 2 is significant at a conventional significance level, and also the adjustment speed in regime 2 is almost 11 times faster than that in regime 1. This suggests that when dividends of previous period are relatively higher than their target, dividends of current period move toward the target rapidly when dividends of previous period are relatively lower than their target.

Table 4.5
Estimation of Threshold VECM

Model	VECM without lag		VECM with 1 lag	
Dependent variable	Δd_t	Δp_t	Δd_t	Δp_t
α_1	-0.017	0.025	**-0.031**	0.009
(s.e.)	(0.011)	(0.023)	(0.011)	(0.023)
μ_1	**-0.015**	0.041	**-0.027**	0.013
(s.e.)	(0.015)	(0.032)	(0.011)	(0.024)
$\Delta d_{1,t-1}$			**0.338**	0.048
(s.e.)			(0.078)	(0.089)
$\Delta p_{1,t-1}$			-0.022	0.098
(s.e.)			(0.025)	(0.055)
α_2	**-0.183**	0.221	**-0.203**	0.031
(s.e.)	(0.031)	(0.288)	(0.043)	(0.439)
μ_2	**-0.199**	0.220	**-0.153**	0.014
(s.e.)	(0.032)	(0.297)	(0.030)	(0.300)
$\Delta d_{2,t-1}$			**0.701**	0.283
(s.e.)			(0.094)	(0.367)
$\Delta p_{2,t-1}$			-0.036	-0.283
(s.e.)			(0.028)	(0.172)
β	0.655		0.598	
γ	-1.125		-0.748	
p_1, p_2	0.818	0.182	0.898	0.102
Likelihood	2,532.125		2,585.071	

The bold number indicates that it is significant statistically at 5% significant level.

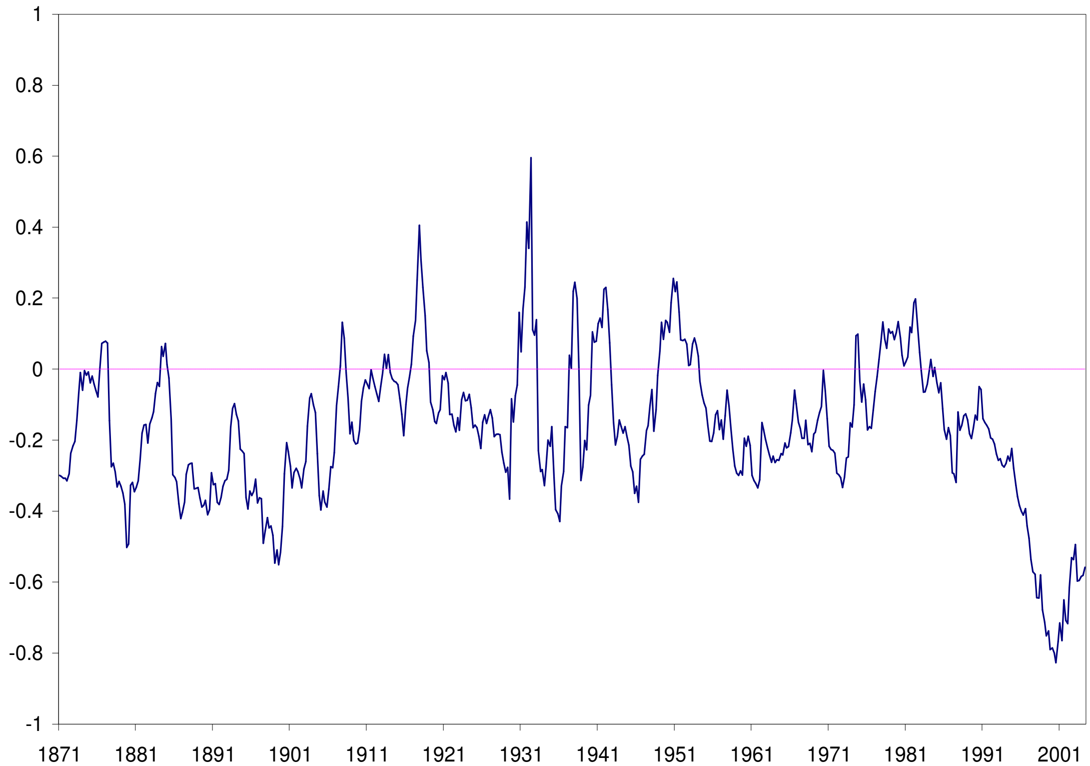

Figure 4.2. The Upper and Lower Regimes

As discussed in above, since adjustment speed coefficients in both regimes are smaller than 0.5, the cost of deviation (ϕ) from the target is still smaller than the cost of adjustment (φ) from the previous dividends. Also we can infer a fact that adjustment cost imposed by managers in lower regime is bigger than that in upper regime (i.e., $\varphi_1 > \varphi_2$) based on the relation between adjustment speed coefficients in equation (4.10) with the parameters in equation (4.8) (i.e., $\alpha_i = \phi/(\phi + \varphi_i)$, where *i*=1, 2). This seems to contradict the fact that managers are reluctant to decrease dividends generally because the investors may take the decrease as bad signal of weak performance of the firm in the future. However, if we consider that the discounted sum of dividend can not exceed to the value of the firm which is equal to stock prices in rational expectation framework, over payout dividend compared to the value of the firm should be adjusted in the future.

Next, we test the hypothesis of asymmetric adjustment of dividends in an extended model which has lagged terms of the changes of dividends and stock prices $(k=1)$. The number of lag is chosen by SIC in maximum lags with 12 under the vector autoregressive model. Generally the estimation results are not much different with those of the model without lag terms. The estimation results of a linear VECM (d_t, p_t) with one lag term are reported in right side of Table 4.3. The adjustment speed coefficient is significant at standard significance level, and also it is not different with that of a linear VECM (d_t, p_t) without lag term. For the lag terms, the coefficient of the lag term of dividends change is significant statistically and it has relatively big value, while the coefficient of the lag term of prices change is not significant statistically. This indicates that dividends change shows some persistent movement in this specification. However, in prices equation the current change of prices does not affect by any independent variables, which supports that past information can not predict the future change of stock prices in efficient market. The test of threshold effect can not be rejected in the model (d_t, p_t) at 5% significance level as shown in bottom of Table 4.4.

Table 4.5 shows the estimation results of a nonlinear VECM with one lag term $(k=1)$. Compared to the nonlinear VECM without lag term, the biggest difference is that the adjustment speed coefficient in lower regime becomes significant statistically. So deviations of dividends from the target in previous period are adjusted by change of current dividends in both regimes. Also the adjustment speeds in both regimes are increased slightly. Compared to the linear VECM with lag term, still only coefficients of lag term of change of dividends at time *t-1* in both regimes are significant statistically, but those of lag term of change of prices are not. Whereas the change of prices are not be influenced by the dividends in both regimes. This still supports the Miller and Modigliani hypothesis in this extended specification. In sum the main findings in the model without lag term do not change in the extended model with lag term such that the hypothesis of regime dependent adjustment cost can not be rejected when stock prices are used for the target. The adjustment cost is bigger than the deviation cost from target, and also the adjustment cost (adjustment speed) in upper regime is much smaller (faster) than that in lower regime.

4.5 Conclusion

This chapter assesses the asymmetric adjustment behavior in aggregate dividends under the nonlinear Garrett and Priestley's (2000) framework that assumes managers minimize the regime dependent adjustment costs associated with being away from their target dividend payout. Real stock prices are used for a proxy for the target. By using the threshold vector error correction model, we find significant evidence of threshold effect in adjustment behavior of aggregate dividends of S&P 500 Index in quarterly data when real stock prices are used a proxy for the target dividend. This result indicates that the speed of adjustment of dividends is different depending on the regime defined the difference between dividends and target in previous period. More specifically, the adjustment speed of dividends in upper-regime, where dividends are relatively higher than the target, is much faster than that in lower-regime, where dividends are relatively lower than the target.

In this chapter we only consider asymmetric adjustment of dividends based on the degree of the difference between dividends and target, but the other form of asymmetric adjustment of dividends also may be plausible. For example, as suggested by Jalilvand and Harris (1984), adjustment speed of dividends might depends on the condition of capital market such interest rates, stock prices etc. We leave this as future works.

CHAPTER V

SUMMARY

This dissertation investigates three different economic issues. The first issue is whether the monetary policy rules can improve forecasting accuracy of inflation. The second one is whether the preference of central bank is symmetry or not. The last issue is whether the behavior of aggregate dividends is asymmetry. Each issue is covered in Chapter II, III and IV, respectively.

The Chapter II begins with the rational expectation model. This model implies that nominal interest rates reflect expectations of inflation, and thus the term structure of interest rates provides information on the future change in inflation. However, the monetary authority manipulates the short-term interest rate in response to the change in the price level, and accordingly the prediction of inflation cannot be separated from the monetary policy.

In this chapter, we undertake an empirical analysis of this linkage using the U.S. monthly data for the period January 1960-December 2004. First, we estimate the long-run Taylor rule, which is composed of the federal funds rate and the 12-month inflation rate. The coefficient of reaction to inflation varies depending on the sample period and across the monetary policy regimes. Second, the prediction of inflation is found to be associated with the Fed's reaction to inflation. The coefficient of the term structure is significant for the sample period when the coefficient of reaction to inflation is close to unity. As the parameter of reaction to inflation increases, the predictive information contained in the term structure becomes weaker. This result explains the previous empirical findings that the predictive information of the term structure varies depending on the sample period. Third, an assessment of the prediction performance regarding future change in inflation is provided using the long-run Taylor rule. The empirical results indicate that the long-run Taylor rule improves forecasting accuracy.

In the Chapter III, we investigate the asymmetric preferences of central bank of Korea under New Keynesian sticky prices forward-looking economy framework. In this end, this chapter adopts the central bank's objective functional form as the linear-exponential function over inflation gap or/and output gap instead of the standard quadratic function.

We derive the monetary policy reaction function, and then estimate the derived policy reaction function during the inflation targeting period: 1998:9-2005:12. With caution we conclude that the AS curve may not be nonlinear form because the cross-product term of inflation gap and output gap is not significant and its sign is wrong. Second, with a linear AS curve the Bank of Korea have an asymmetric preference over inflation gap but not output gap. Third, since the sign of the preference parameter is positive, it appears that the Bank of Korea has the more weight on positive deviations of inflation from target than on negative ones, which supports the hypothesis of precautionary demand for inflation.

The Chapter IV assesses the asymmetric adjustment behavior in aggregate dividends. In this end, we derived the nonlinear dividend adjustment model under the hypothesis that managers minimize the regime dependent adjustment costs associated with being away from their target dividend payout.

By using the threshold vector error correction model, we find significant evidence of threshold effect in adjustment behavior of aggregate dividends of S&P 500 Index in quarterly data when stock prices are used a proxy for the target dividend. This indicates that the adjustment cost is regime-dependent. We also find that when dividends are relatively higher than target in previous period, the adjustment cost (speed) of dividends is smaller (faster) than that of dividends when they are relatively lower.

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APPENDIX

ASYMMETRIC PREFERENCE

UNDER PARTIAL ADJUSTMENT OF INTEREST RATE

In this appendix we test the asymmetric preferences of central bank where its objective function is different with that in Section 3.2.2, i.e., $(i_t - i^*)^2$ replaces $(i_t - i_{t-1})^2$ in the period loss function for interest rate stabilization, which is used in Surcio (2003b) and Karagedikli and Lees (2004), etc. Therefore, the loss function has the following equation:

$$V_t = \left[\frac{\exp\{\varphi(\pi_t - \pi^*)\} - \varphi(\pi_t - \pi^*) - 1}{\varphi^2}\right] + \theta\left[\frac{\exp(\phi\tilde{y}_t) - \phi\tilde{y}_t - 1}{\phi^2}\right] + \frac{\kappa}{2}(i_t - i^*)^2, \qquad \text{(A.1)}$$

where the parameter κ is strictly positive and govern the relative weight that central bank places on interest rate stabilization relative to inflation stabilization.

With the same economy framework used in Section 3.2.1, the first order condition (FOC) for minimizing the loss of central bank can be derived as

$$-E_{t-1}\left[\frac{\exp\{\varphi(\pi_t - \pi^*)\} - 1}{\varphi}\right]\frac{\alpha\beta}{(1-\beta\psi\tilde{y}_t)^2} - E_{t-1}\left[\frac{\exp(\phi y_t) - 1}{\phi}\right]\theta\alpha + \kappa(i_t - i^*) = 0. \quad \text{(A.2)}$$

By taking the first order Taylor series expansion at $\varphi = \phi = \psi = 0$, then equation (A.2) can be written as

$$-\alpha\beta E_{t-1}(\pi_t - \pi^*) - \theta\alpha E_{t-1}\tilde{y}_t - \frac{\varphi\alpha\beta}{2}E_{t-1}(\pi_t - \pi^*)^2 +$$

$$-\frac{\phi\theta\alpha}{2}E_{t-1}\tilde{y}_t^2 - 2\psi\alpha\beta^2 E_{t-1}[(\pi_t - \pi^*)\tilde{y}_t] + \kappa(i_t - i^*) + \varepsilon_t = 0, \qquad \text{(A.3)}$$

where ε_t represents the higher order of the Taylor series expansion. Solving for i_t and the expected inflation and output gap are replaced by actual values, then the nonlinear monetary policy reaction function can be derived. However, it is generally believed that central banks have a tendency to smooth changes in interest rates. Following the literature such as Clarida et al. (1998), the actual interest rate is assumed to partially adjusts to the target such as:

$$i_t = (1-\rho)i_t^* + \rho i_{t-1}. \tag{A.4}$$

Combining this partial adjustment process with equation (A.3), then the empirical interest rate reaction function can be written as follows:

$$i_t = (1-\rho)\{c_1 + c_2\tilde{\pi}_t + c_3\tilde{y}_t + c_4\tilde{\pi}_t^2 + c_5\tilde{y}_t^2 + c_6(\tilde{\pi}_t\tilde{y}_t)\} + \rho i_{t-1} + u_t, \tag{A.5}$$

which is linear in the coefficients. Also we get the following coefficients and error term conditions related with equation (A.3) and (A.5):

$$c_0 \equiv i^*, \quad c_1 \equiv \frac{\alpha\beta}{\kappa}, \quad c_2 \equiv \frac{\theta\alpha}{\kappa}, \quad c_3 \equiv \frac{\varphi\alpha\beta}{2\kappa}, \quad c_4 \equiv \frac{\phi\theta\alpha}{2\kappa}, \quad c_5 \equiv \frac{2\psi\alpha\beta^2}{\kappa}, \text{ and}$$

$$\frac{u_t}{(1-\rho)} \equiv -\left\{\begin{matrix} c_1(\tilde{\pi}_t - E_{t-1}\tilde{\pi}_t) + c_2(\tilde{y}_t - E_{t-1}\tilde{y}_t) + c_3(\tilde{\pi}_t^2 - E_{t-1}\tilde{\pi}_t^2)] \\ c_4(\tilde{y}_t^2 - E_{t-1}\tilde{y}_t^2) + c_5[\tilde{\pi}_t\tilde{y}_t - E_{t-1}(\tilde{\pi}_t\tilde{y}_t)] \end{matrix}\right\} + \frac{\varepsilon_t}{\kappa}.$$

The error term is a linear combination of forecast errors and thus orthogonal to any variable in the information set available at t-1. Therefore, based on this orthogonality condition, coefficients in equation (A.5) can be estimated by the generalized method of moments (GMM). From the reduced coefficients, we can recover the asymmetric preferences over inflation gap and output gap such as $\varphi = 2c_3/c_1$, and $\phi = 2c_4/c_2$.

The central bank's reaction function, i.e., equation (A.5), is estimated by the GMM using the same instruments in Section 3.3.2. Table A.1 reports the results of estimating reaction functions. Over various specifications the estimates of interest rate smoothing coefficients are all significant and are ranged from 0.848 to 0.923, which are greater than those for Australia, New Zealand, US and European Area documented by Karagedikli and Lees (2004) and Surcio (2003b) using the similar framework. This means that the Bank of Korea has changed the interest rate toward the target rate very slowly compared to other central banks. The constant estimates, which represent the target interest rate, in nonlinear reaction functions are very close to a value of 4.032 in linear case.

In here, we look at the results of estimating the nonlinear reaction function with square terms of inflation gap and output gap shown in Panel D. When the cross term is not considered, square of inflation gap but output gap is significant at 1% significance level. It seems to support only asymmetric preference over inflation. However, the cross term is included, square terms of both inflation gap and output gap are significant but coefficient of output gap become insignificant. We estimate asymmetric parameters and test their significance, and the estimation results are reported in Table A.2. When only inflation asymmetry or output asymmetry is considered, they are significant and have expected sign suggested by Cukierman and Muscatelli (2003). However, only inflation asymmetry is significant when preference asymmetries over inflation gap and output gap are considered. This result is consistent with that in Section 3.3.2. But its size is much bigger than that (i.e., 0.483) reported in there.

Table A.1
Reduced Form Estimates of Reaction Function- Partial Adjustment

Model	ρ	c_1	c_2	c_3	c_4	c_5	c_6	J	$\overline{R}^2$
Panel A: Linear reaction function									
(1)	0.888**	4.032**	0.343**	0.233**				0.131	0.980
	(0.008)	(0.073)	(0.103)	(0.035)					
Panel B: Nonlinear reaction function with inflation gap only									
(2)	0.855**	3.922**	0.746**	0.193**	0.429**			0.118	0.978
	(0.011)	(0.083)	(0.145)	(0.022)	(0.069)				
(3)	0.873**	3.940**	0.780**	0.158**	0.329**		-0.062*	0.120	0.982
	(0.010)	(0.093)	(0.162)	(0.035)	(0.065)		(0.030)		
Panel C: Nonlinear reaction function with output gap only									
(4)	0.923**	4.037**	0.047	0.352**		-0.019**		0.103	0.983
	(0.011)	(0.083)	(0.135)	(0.062)		(0.006)			
(5)	0.922**	4.046**	-0.002	0.375**		-0.021**	0.023	0.102	0.983
	(0.011)	(0.088)	(0.136)	(0.081)		(0.007)	(0.036)		
Panel D: Nonlinear reaction function with inflation gap and output gap									
(6)	0.848**	3.914**	0.756**	0.185**	0.434**	0.002		0.116	0.976
	(0.017)	(0.082)	(0.143)	(0.027)	(0.068)	(0.003)			
(7)	0.849**	3.819**	1.148**	0.023	0.444**	0.013*	-0.124*	0.102	0.975
	(0.022)	(0.136)	(0.270)	(0.070)	(0.109)	(0.005)	(0.052)		

Estimation model is as follows: $i_t = (1-\rho)\{c_1 + c_2\tilde{\pi}_t + c_3\tilde{y}_t + c_4\tilde{\pi}_t^2 + c_5\tilde{y}_t^2 + c_6(\tilde{\pi}_t\tilde{y}_t)\} + \rho i_{t-1} + \mu_t$. The estimation period is from 1998:9 to 2005:12. Output gap is obtained from percentage change of actual and detrending industrial production index with quadratic trend equation for the sample period 1998:1-2005:12. The instrument sets include constant and five lags of core inflation, output gap, long-short interest rate spread and call rate. The superscript **, * and + denote the rejection of the null hypothesis that the true coefficient is zero at the 1 percent, 5 percent and 10 percent significance levels, respectively.

Table A.2
Estimates of the Asymmetric Policy Preferences- Partial Adjustment

Model	Inflation gap (φ)	Output gap (γ)	J	$\overline{R}^2$
Panel A: Inflation gap only				
(2)	0.783**	-	0.122	0.983
	(0.106)			
(3)	0.844**	-	0.120	0.982
	(0.099)			
Panel B: Output gap only				
(4)	-	-0.108**	0.103	0.983
		(0.021)		
(5)	-	-0.110**	0.102	0.983
		(0.020)		
Panel C: Inflation gap and output gap				
(6)	1.148	0.022	-	-
	(-)	(-)		
(7)	0.744**	1.117	0.102	0.975
	(0.134)	(3.700)		

The estimation period is from 1998:9 to 2005:12. Output gap is obtained from percentage change of actual and detrending industrial production index with quadratic trend equation for the sample period 1998:1-2005:12. The instrument sets include constant and five lags of CPI inflation, output gap, long-short interest rate spread and call rate. The superscript **, * and + denote the rejection of the null hypothesis that the true coefficient is zero at the 1 pecent, 5 percent and 10 percent significance levels, respectively.

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