Discussion Paper Series
No. 2007
August 2020

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August, 2020

Abstract

We examine the role of the Fed's preference in the understanding of inflation rate and unemployment rate evolution using US data over the period of 1960–2017. Facing the evidence of instability in a constant-coefficient regression, we run a nonparametric regression, and find that the Fed's preference parameters have moved, implying that its preference can be represented by the asymmetric preference model putting more weights on high unemployment rate approximately before the era of Volcker's chairmanship and by the inflation targeting model during the 1980s and 1990s. The Fed's preferences again seem concerned about higher unemployment after the Global Financial Crisis.

KEYWORDS
asymmetric preference, inflation, monetary policy, time-varying parameter, unemployment, nonparametric regression

JEL CLASSIFICATIONS
E31, E52, E61

1. INTRODUCTION

The dynamics of the rates of inflation and unemployment in the US have attracted the attention of many economists. Studies by Kydland and Prescott (1977), Barro and Gordon (1983), and Cukierman and Gerlach (2003) are among the numerous attempts to explain the behavior of both the inflation and unemployment rates based on a central banker's preference and rational expectation, whereby equilibrium often results in the problem of time-inconsistency. Additionally, the empirical validity of such studies is examined to test the reliability of the interpretation based on the time-inconsistency problem with the data of the inflation rate and unemployment rate. Ireland (1999) shows that the time-inconsistency problem in the Barro–Gordon model explains the long-run cointegration relation between the two variables. Ruge-Murcia (2003) compares the Barro–Gordon model with a version of the Cukierman–Gerlach model (the model with asymmetric loss function) and demonstrates that the asymmetric loss function similar to that of Cukierman and Gerlach (2003) is more compatible with the data.

Following Ireland (1999) and Ruge-Murcia (2003), we also investigate which type of central banker's preference is supported by the data on the rates of inflation and unemployment. In the process, we explain the fluctuation in the inflation and unemployment rates, by allowing a time-variation in the Fed's preference parameters which govern the degree of asymmetry for positive and negative deviations in unemployment from the target in the preference, the relative weight between the inflation and unemployment components in the preference, and the unemployment rate target. Lakdawala (2016) argues that the Fed's preference parameters evolve gradually because the composition of the Federal Open Market Committee (FOMC) varies over time along with the political pressure on the Fed. Hence, we examine how the variation in these parameters results in different relationships between the inflation rate and unemployment rate, derived from the first-order condition of the Fed's optimization problem.

Our approach is similar to Lakdawala (2016) in the sense that we allow variation in

the Fed's preference parameters over time. However, we employ the nonparametric approach developed by Andrews (1991a) to estimate the time-varying relationship. Although the Kalman filtering approach used in Lakdawala (2016) requires a specific process for parameters to be estimated, our approach does not impose any process or functional form for time-varying parameters *a priori*. In addition, the first-order condition in our study is derived from an asymmetric loss function which nests a quadratic loss function, assumed in Lakdawala (2016), a special case.

The remainder of the study is organized as follows. Section 2 includes a brief survey of related literature. Section 3 presents the Fed's preference (or loss function), nesting the Barro–Gordon model, the Cukierman–Gerlach model, or the inflation targeting model as a special case.[1] Similar to Ruge-Murcia (2003), a relationship between the inflation rate, unemployment rate, and the volatility of the unemployment rate is derived from the first-order condition of the Fed. Section 4 provides a brief description of the data used in this study. In Section 5, we provide the findings of the constant-coefficient regression, presenting evidence that its coefficients are unstable. Hence, we employ a nonparametric time-varying regression to illustrate the variation of the coefficients over time. We also relate the variation of these coefficients to the variation of the Fed's preference. The results from the constant-coefficient regression support the Cukierman–Gerlach model for the Fed's preference and provide completely different interpretations from the nonparametric regression. The results of the nonparametric regression suggest that the Cukierman and Gerlach (2003) model can describe the Fed's preference before the Volcker period and after the 2008 global financial crisis (hereafter GFC), and the inflation-targeting model during the 1980s and 1990s. Our results also

[1] Surico (2007) examines the Fed's loss function, which assumes an asymmetric reaction to both inflation and output gap, and finds that the parameter related to the asymmetric reaction to inflation is never significant. Hence, this study follows Ruge-Murcia (2003) to investigate the Fed's loss function with an asymmetric reaction to unemployment and a symmetric reaction to inflation.

imply that changes in the monetary policy rule since Volcker's era, reported in many studies, may be related to movements in the Fed's preferences. We further provide evidence of the evolution of the Fed's preferences during the 2000s and 2010s, which has been discussed by only a few studies until now. Section 6 provides the concluding remarks.

2. LITERATURE REVIEW

Many studies have examined the role of a central bank in understanding the dynamics of macroeconomic performance. Studies including Kydland and Prescott (1977), Barro and Gordon (1983), Ireland (1999), Cukierman and Gerlach (2003) and Ruge-Murcia (2003), which are cited in the Introduction, attempt to understand the dynamics of the inflation rate and unemployment rate based on a central banker's preference and rational expectation. However, Sargent (1999) acknowledges the difficulty of explaining the rise and fall of the inflation rate using one type of time-inconsistency model with the rational expectation hypothesis.

Three approaches are proposed to overcome this challenge. First, Sargent (1999) emphasizes a continual learning process that can be interpreted as the variation in the Fed's view of the behavior of the private sector. Second, Lakdawala (2016) introduces the changes in central bankers' preferences to the time-inconsistency model. Finally, Sims and Zha (2006) argue that the time-varying volatility of macroeconomic shocks is critical in explaining the behavior of macroeconomic variables in the US.[2]

Meanwhile, numerous studies have reported that monetary policy rules have varied over time. Limited examples of such works are Clarida, Gali, and Gertler (2000), Dolado, Pedrero, and Ruge-Murcia (2004), Surico (2007, 2008), Bae, Kim, and Kim (2012), Bianchi (2013), Fiodendji (2013), and Best and Hur (2019). Among them, studies such as Dolado et al.

[2] Best and Hur (2019) examine the contributions of these three sources in understanding the time-variation of US monetary policy between 1960 and 2007.

(2004) and Surico (2007, 2008) estimate US monetary policy reaction functions based on non-quadratic loss functions and suggest that the Fed's preferences showed a shift between the pre-Volcker and Volcker–Greenspan eras. In addition, the study by Fiodendji (2013) estimates the Canadian monetary policy reaction function and asymmetric preference parameters for various sub-periods and regimes. It further provides evidence that since 1991, the preferences of the Canadian monetary authority have changed. While it could be hypothesized that variations in the monetary policy rule could be related to or caused by shifts in the central banker's preference, Debortoli and Nunes (2014) show that changes in the parameters of a monetary policy rule do not necessarily correspond to changes in a monetary authority's preference.

Hence, similar to Dennis (2006) and Lakdawala (2016), as mentioned in the second approach above, it is worth explaining the behavior of the rates of inflation and unemployment for an extended sample period based only on variations in the Fed's preference. Moreover, it is worth comparing our results with the prior studies examining the time-variation of US monetary policy rules to test whether changes in them are related to changes in the policymakers' preferences. In doing so, however, our approach is different from previous studies such as Fiodendji (2013) and Lakdawala (2016) as we do not assume a preconceived number of breaks or known structural break dates. We consider the possibility that the relationships among the variables of inflation and unemployment are not permanently fixed, and the coefficients in the reduced-form regression evolve gradually and smoothly because the Fed's preference parameters vary gradually over time.

3. AN ECONOMIC MODEL WITH TIME-VARYING PARAMETERS

The preference of the Fed (or the loss function) is represented by the following function:

$$L(\pi_t, u_t; \pi_t^*, \kappa_t, \lambda_t, \gamma_t) = \frac{1}{2}(\pi_t - \pi_t^*)^2$$

$$+\frac{\lambda_t}{\gamma_t^2}\left[exp\left(\gamma_t(u_t-\kappa_t E_{t-1}(u_t^n))\right)-\gamma_t(u_t-\kappa_t E_{t-1}(u_t^n))-1\right] \quad (1)$$

where π_t denotes the inflation rate, u_t denotes the unemployment rate, u_t^n denotes the natural rate of unemployment, and π_t^* denotes the implicit target (or desirable) inflation rate set by the Fed at time $t-1$. Ruge-Murcia (2003) uses equation (1) to test whether the Barro–Gordon quadratic preference or the Cukierman–Gerlach asymmetric preference is supported by data on the inflation rate and unemployment rate. κ_t, λ_t, and γ_t are the Fed's preference parameters, which are assumed to be given at time $t-1$ by the composition of FOMC members, or political pressure in a central banker's mind, and known to the public through full understanding of the central banker's problem.

κ_t is related to the implicit target of the unemployment rate. When the implicit target of the Fed for the unemployment rate is lower than the expected natural unemployment rate (that is, $\kappa_t E_{t-1}(u_t^n) < E_{t-1}(u_t^n)$ or $0 < \kappa_t < 1$), then it tends to generate inflation bias, as in the model of Barro and Gordon (1983). γ_t governs the degree of asymmetry for positive and negative unemployment deviations from the target in the preference. When $\gamma_t > 0$, a central banker places more weight on the rate of unemployment during a recession as compared to a period of a boom with regard to the loss function. As $\gamma_t \to 0$, the unemployment component in the preference takes a quadratic form.[3] When $\gamma_t \to 0$ with $0 < \kappa_t < 1$, the loss function in equation (1) collapses to the quadratic loss function in the Barro–Gordon model that was tested by Ireland (1999). Finally, $\lambda_t \geq 0$ determines the relative weight between the inflation and unemployment components in the preference. A smaller value for λ_t indicates greater importance for the inflation component in the loss function, and $\lambda_t = 0$ means strict inflation targeting. Hence, the representation of the Fed's preference in equation (1) nests the

[3] This can be verified by applying L'Hôpital's rule twice.

Barro–Gordon model ($\gamma_t \to 0$ and $\lambda_t > 0$, with $0 < \kappa_t < 1$), the Cukierman–Gerlach type model ($\gamma_t > 0$ and $\lambda_t > 0$ with $\kappa_t = 1$), or the inflation targeting ($\lambda_t = 0$ with $\gamma_t \neq 0$, or $\gamma_t \to 0$ with $\kappa_t = 1$) as a special case.[4]

In addition to a central banker's preference, other parts of the model are standard in the literature. We assume the following expectation-augmented Phillips curve:

$$u_t = u_t^n - \alpha\big(\pi_t - E_{t-1}(\pi_t)\big) + v_t, \tag{2}$$

where α is assumed to be positive, implying a non-vertical Phillips curve, and v_t denotes a supply-side disturbance. $E_{t-1}(\pi_t)$ denotes the expectation of the inflation rate conditional on all the data available at time $t-1$. We assume that the public has the same set of information as the Fed, and they form rational expectations along with the central banker.[5]

The natural rate of unemployment (u_t^n) evolves as a stationary process.[6] The parameters in the expectation-augmented Phillips curve and the process of u_t^n could vary over time because of technological developments, demographic changes, or other changes in the economic environment. However, we assume that the parameters are constant in order to see how much we can explain the behavior of the rates of inflation and unemployment, based only on variations in the central bank's preference parameters.

Finally, we assume that the Fed can affect the inflation rate as follows:

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[4] Barro and Gordon (1983) suggest that the solution $\pi_t = \pi_t^*$ could be achieved when the policymaker sets zero weight to unemployment (λ_t=0) or considers the natural rate of unemployment as optimal ($\kappa_t = 1$). Svensson (1999) names the loss function with the inflation component only (λ_t=0) as "strict inflation targeting." We term all these cases ($\lambda_t = 0$ with $\gamma_t \neq 0$, or $\gamma_t \to 0$ with $\kappa_t = 1$) "the inflation targeting."

[5] This assumption might be strong, however, there is no consensus about the level of difference in the information set of the public and the Fed. The Fed and mass media play an important role in reducing the gap in the information set of the Fed and the public. For example, the Fed regularly publishes various forms of information (data, reports, predictions, etc.), and the mass media readily provides the public with an interpretation of this published information. Hence, we assume that the public and the central bank have the same information set, which also simplifies the analysis of this study.

[6] Ireland (1999) assumes a nonstationary process for u_t^n, and Ruge-Murcia (2003) assumes ARIMA (2,0,2) or ARIMA (1,1,2) for u_t^n. However, various tests in Table 2 later indicate that the rate of unemployment follows a stationary process. Hence, we assume that u_t^n also follows a stationary process. In the appendix, we also consider the possibility that the unemployment rate follows a nonstationary process.

$$\pi_t = i_t + \eta_t, \tag{3}$$

where i_t denotes the policy instrument of a central banker and η_t denotes a control error. The specification in equation (3) states that while the Fed can affect the rate of inflation, its control is not perfect.

A central banker sets a policy instrument to minimize the loss function in equation (1) subject to the expectation-augmented Phillips curve, equation (3), and the anticipation of the private agents formed at time $t-1$ regarding the behavior of the central banker at time t.[7] The first-order condition can be written as

$$E_{t-1}(\pi_t) = \pi_t^* + \frac{\alpha\lambda_t}{\gamma_t} E_{t-1}\left(\exp\left(\gamma_t\left(u_t - \kappa_t E_{t-1}(u_t^n)\right)\right) - 1\right). \tag{4}$$

As explained in Ruge-Murcia (2003), the assumptions of the rational expectation hypothesis and the normal distribution for the rate of unemployment imply that equation (4) can be rewritten as

$$\pi_t = \pi_t^* + \frac{\alpha\lambda_t}{\gamma_t}(\exp(\gamma_t(1-\kappa_t)E_{t-1}(u_t^n) + \frac{\gamma_t^2\sigma_{u,t}^2}{2}) - 1) + \eta_t, \tag{5}$$

where $\sigma_{u,t}^2$ is the conditional variance of u_t. Hence, the first-order Taylor approximation and the replacement of $E_{t-1}(u_t^n)$ with $E_{t-1}(u_t)$ provide the following reduced-form relationship among π_t, $E_{t-1}(u_t)$, and $\sigma_{u,t}^2$:

$$\pi_t = a_t + b_t E_{t-1}(u_t) + c_t\sigma_{u,t}^2 + e_t, \tag{6}$$

where e_t is a disturbance in the reduced form. π_t^* can vary over time but affects only a_t in equation (6), while the Fed's preference parameters affect b_t and c_t. In the estimation, the time-variation of π_t^* will be reflected in the disturbance term of the regression. While we cannot identify the Fed's preference parameters (κ_t, λ_t, and γ_t) separately from the reduced form regression in equation (6), preference parameters have relationships with b_t and c_t

[7] The structural shocks in the model are assumed to be serially uncorrelated and normally distributed with mean zero and heteroscedastic variance, as in Ruge-Murcia (2003).

depending on the central banker's preference. Since $b_t = \alpha\lambda_t(1-\kappa_t)$ and $c_t = \frac{\alpha\lambda_t\gamma_t}{2}$, we have three possible interpretations for the estimation results from regression (6).

When the Barro–Gordon model ($\gamma_t \to 0$ and $\lambda_t > 0$, with $0 < \kappa_t < 1$) can represent the Fed's preference, then b_t will be positive and $c_t = 0$. If the Fed's preference can be represented by the Cukierman–Gerlach model ($\gamma_t > 0$ and $\lambda_t > 0$, with $\kappa_t = 1$), then $b_t = 0$ and c_t will be positive.[8] Finally, if the inflation targeting describes the Fed's preference ($\lambda_t = 0$ with $\gamma_t \neq 0$, or $\gamma_t \to 0$ with $\kappa_t = 1$), then $b_t = 0$ and $c_t = 0$. We investigate which implication is supported by data.

In this exercise, we assume that the Fed's preference parameters vary over time. While we have no direct evidence of variations in the preference parameters, various studies report that the monetary policy function of the Fed has changed over time. For example, Clarida et al. (2000) demonstrate that the Fed's estimated monetary policy rule implies that since 1979, it has placed more weight on the inflation component in the loss function. Bae et al. (2012) argue that the evolution of the monetary policy rule of the Fed approximately coincides with the terms of its various chairpersons. Bianchi (2013) and Best and Hur (2019) contend that a change in the monetary policy rule is one of the important factors in understanding the macroeconomic dynamics of the US. Debortoli and Nunes (2014) demonstrate that changes in the parameters of the monetary policy rule are not necessarily linked to changes in the Fed's preference. Thus, we attempt to examine whether the findings described in existing studies are related to changes in the Fed's preference parameters by comparing them with our results. The current study intends to supplement the literature in this regard.

[8] When $\kappa_t = 1$, one cannot rule out the possibility that $\gamma_t < 0$ and $\lambda_t < 0$, which also results in $b_t = 0$, and a positive c_t. Hence, we do not claim that the condition we mentioned is necessary and sufficient to distinguish the Cukierman–Gerlach model from the Barro–Gordon model. Thus, the abovementioned condition should be interpreted as suggestive information, rather than a necessary condition.

4. DATA

Quarterly US data on inflation and unemployment are used for this study's analysis. The percentage change in the quarterly GDP implicit price deflator and the quarterly average of the monthly civilian unemployment rate are taken for the construction of the inflation rate and the unemployment rate, respectively. These seasonally adjusted data are obtained from the website of the Federal Reserve Bank of St. Louis.[9] The sample starts in the first quarter of 1960 and ends in the third quarter of 2017. The descriptive statistics of the variables are summarized in Table 1. Table 1 provides information on the mean, maximum, minimum, standard deviation, and the number of observations for the rates of inflation and unemployment.

TABLE 1 Descriptive statistics

	Unemployment rate	Inflation rate
Mean	6.0489	0.8183
Maximum	10.6667	2.9420
Minimum	3.4000	-0.1453
Standard deviation	1.5799	0.5887
Number of observations	231	231

Notes: The inflation rate is calculated as the percentage change in the GDP implicit price deflator from quarter-to-quarter, and the unemployment rate is measured as the quarterly average of the monthly civilian unemployment rate.

5. EMPIRICAL ANALYSIS

Table 2 reports the results of various unit root tests. The results are a little sensitive depending on the method of the test. For example, the augmented Dickey–Fuller (ADF) test rejects the unit root null hypothesis for the unemployment rate but does not reject for the inflation rate at the 5% significance level. However, the test results are different when applying the Phillips–Perron test. In this case, the unit root null hypothesis is rejected for the unemployment rate at

[9] The web address is https://fred.stlouisfed.org/.

the 10% significance level, while the null hypothesis is rejected for the inflation rate at the 5% level. Since these indecisive results may be due to the lack of the power in the ADF test and the Phillips–Perron test, we further conducted the Elliott–Rothenberg–Stock (1996) DF-GLS unit root test, which is known to have superior power, particularly with a small sample size. The unit root null can be rejected for all variables at the 5% level using the DF-GLS test. Moreover, we conducted the Kwiatkowski–Phillips–Schmidt–Shin test, and the null hypothesis of stationarity is not rejected for any of the variables. Hence, based on these findings, we assume that both the unemployment rate and inflation rate are stationary.[10]

TABLE 2 Unit root test results

	Unemployment rate	Inflation rate
Augmented Dickey–Fuller	-3.1654**	-2.3788
Phillips and Perron	-2.6700*	-3.1802**
DF-GLS	-1.9575**	-2.6129***
Kwiatkowski–Phillips–Schmidt–Shin	0.2157	0.3206

Notes: When conducting a unit root test, an intercept is included in a test equation. A lag length is selected by the Akaike information criterion for the augmented Dickey–Fuller and Elliott–Rothenberg–Stock DF-GLS tests. Andrews' (1991b) bandwidth selection is used for the Phillips–Perron test and the Kwiatkowski–Phillips–Schmidt–Shin test. *, **, and *** denote that the null hypothesis is rejected at the 10%, 5%, and 1% levels, respectively.

5.1 Constant coefficient regression

Table 3 shows the results of constant-coefficient regression for equation (6) when the Fed's preference parameters, κ_t, λ_t, and γ_t, are assumed to be constant. The second row shows the result when $E_{t-1}(u_t)$ is replaced with $u_t + \varsigma_t$ (actual unemployment rate and forecast error)

[10] Similar to Ireland (1999), we examine the relationship between the unemployment rate and inflation rate when both are nonstationary and cointegrated, and the results are in the appendix.

under the rational expectation hypothesis.[11] Further, $\sigma_{u,t}^2$ is the realized variance of the unemployment rate through the use of monthly actual unemployment rates during each quarter. If ς_t (forecast error) is uncorrelated with e_t (disturbance in the reduced form) under the rational expectation hypothesis, the coefficients in equation (6) can be consistently estimated. As shown in the second row, b (the coefficient of $E_{t-1}(u_t)$) is insignificant, and c (the coefficient of the volatility of u_t) is significantly positive at the 10% level. These results support the Cukierman–Gerlach-type asymmetric model. When $E_{t-1}(u_t)$ and $\sigma_{u,t}^2$ are constructed from the ARIMA (2,0,2) with GARCH (1,1) assumption for the unemployment rate, as in Ruge-Murcia (2003), b is insignificant while c is significantly positive at the 1% level.[12] This result is consistent with Ruge-Murcia (2003). In our longer sample period, the data support the Cukierman–Gerlach-type asymmetric model, while adjusted $\bar{R}^2$s are low ranging, between 0.034 and 0.055. These results together suggest the Cukierman–Gerlach-type model for the Fed's preference when performing the constant-coefficient regressions.

TABLE 3 Constant coefficient regression results

	a	b	c	adjusted $\bar{R}^2$
$E_{t-1}(u_t) = u_t + \varsigma_t$ and $\sigma_{u,t}^2$ is the realized variance of u_t	0.6943** (0.3274)	0.0119 (0.0539)	3.4660* (2.0545)	0.0344
$E_{t-1}(u_t)$ and $\sigma_{u,t}^2$ are from the ARIMA (2,0,2) model with GARCH (1,1) for u_t	0.9585** (0.3937)	-0.0490 (0.0623)	2.8249*** (0.9422)	0.0554

Notes: This table displays the regression results for $\pi_t = a + bE_{t-1}(u_t) + c\sigma_{u,t}^2 + e_t$. The sample period is 1960: I–2017: III. The numbers in parentheses are Newey–West standard errors.

[11] ς_t denotes a forecast error.

[12] Results of the constant-coefficient regression robustly support the Cukierman–Gerlach-type asymmetric model even when ARMA processes other than ARIMA (2,0,2) are used in the analysis for the unemployment rate. The results are available upon request.

If a version of the Cukierman–Gerlach model is a good description of the Fed's preference for the entire sample period, however, it is a little difficult to justify the change in the stance of the Fed in lowering and stabilizing inflation since Volcker's era. Figure 1 shows that despite the elevated level and considerable volatility of the rate of unemployment, the US inflation trend started falling in the early 1980s. Based on the results in Table 3, a possible interpretation of the movements of the inflation rate since the 1980s would be that the marginal benefit of unexpected inflation has shrunk as the volatility of the unemployment rate has exogenously decreased. Hence, the inflation has been stabilized as the reduced volatility of the unemployment rate has contributed to a lower inflation bias rather than the Fed's initiation to reduce the inflation rate. According to Table 3, this interpretation is similar to Sims and Zha (2006) in the sense that the understanding of the macroeconomic dynamics of the US requires an influential role of time-varying volatilities rather than changes in the Fed's stance.

(Figure 1 here)

However, the findings of the constant-coefficient regressions could be misleading because variations in the monetary policy rule and/or changes in the Phillips curve relationship reported in various studies may imply shifts in the Fed's preference. This conjecture suggests that time-varying coefficients in the reduced form equation (6) should be tested. We conduct structural break tests for the coefficients in the constant-coefficient regression to determine whether they are unstable during the sample period. The Bai and Perron (1998) test is conducted for the constant-coefficient regression, and the results are presented in Table 4. As shown in Table 4, the number of breaks and estimated breakpoints are quite sensitive to test statistics (e.g., F-statistic, UDmax statistic, WDmax statistic, Schwarz criterion, or LWZ criterion) and the assumption on the maximum number of breaks. However, all the results

strongly indicate that the coefficients in the constant-coefficient regression are unstable. Irrespective of the assumption on the maximum number of breaks, the estimated breakpoints can be found in the 1960s, 1970s, 1980s, 1990s, and 2000s, which may imply gradual changes of the coefficients in the regression.

TABLE 4 Multiple breakpoint tests

Methods	Selection	The number of breaks	Break dates	The number of breaks	Break dates
		Maximum # of breaks : 6		Maximum # of breaks : 8	
1 to M globally determined breaks	Sequential F-statistic determined breaks	6	1967Q3, 1973Q3, 1983Q1, 1991Q4, 2000Q1, 2009Q4	8	1967Q3, 1973Q3, 1982Q4, 1988Q2, 1993Q4, 1999Q2, 2004Q4, 2010Q2
	Significant F-statistic largest breaks	6	1967Q3, 1973Q3, 1983Q1, 1991Q4, 2000Q1, 2009Q4	8	1967Q3, 1973Q3, 1982Q4, 1988Q2, 1993Q4, 1999Q2, 2004Q4, 2010Q2
	UDmax determined breaks	3	1967Q3, 1973Q3, 1983Q1	8	1967Q3, 1973Q3, 1982Q4, 1988Q2, 1993Q4, 1999Q2, 2004Q4, 2010Q2
	WDmax determined breaks	3	1967Q3, 1973Q3, 1983Q1	8	1967Q3, 1973Q3, 1982Q4, 1988Q2, 1993Q4, 1999Q2, 2004Q4, 2010Q2
L + 1 vs. L globally determined breaks	Sequential F-statistic determined breaks	3	1967Q3, 1973Q3, 1983Q1	3	1967Q3, 1973Q3, 1983Q1
	Significant F-statistic largest breaks	6	1967Q3, 1973Q3, 1983Q1, 1991Q4, 2000Q1, 2009Q4	6	1967Q3, 1973Q3, 1983Q1, 1991Q4, 2000Q1, 2009Q4
Global information criteria for 0 to M globally determined breaks	Schwarz criterion selected breaks	4	1967Q3, 1973Q3, 1983Q1, 1991Q4	4	1967Q3, 1973Q3, 1983Q1, 1991Q4
	LWZ criterion selected breaks	3	1967Q3, 1973Q3, 1983Q1	3	1967Q3, 1973Q3, 1983Q1

Notes: Breakpoints are determined through tests of globally determined breaks by Bai–Perron (1998). We obtain the test statistics by using the heteroscedasticity and autocorrelation consistent covariance. The maximum number of breaks is either six or eight. The number of breaks and estimated break dates are determined according to each methodology and reported in the third (or fifth) and fourth (or sixth) columns, respectively. The obtained results are based on the case when the trimming percentage of the sample is set at 10%, and the significance level is 0.05.

5.2 Nonparametric regression

Considering the evidence presented in Table 4, we run the nonparametric regression to capture the time-varying property of the coefficients in equation (6). Assume that the coefficients vary smoothly, such that $b_t = b\left(\frac{t}{T}\right)$ and $c_t = c\left(\frac{t}{T}\right)$, where $b(\cdot)$ and $c(\cdot)$ are smooth functions defined on [0, 1], and T is the sample size. Under the assumption that $b(\cdot)$ and $c(\cdot)$ are sufficiently smooth functions that can be approximated with a series of polynomials and/or trigonometric functions, the reduced form regression in equation (6) can be written as follows:

$$\begin{aligned}
\pi_t &= a_t + b_t E_{t-1}(u_t) + c_t \sigma_{u,t}^2 + e_t \\
&= a\left(\frac{t}{T}\right) + b\left(\frac{t}{T}\right) E_{t-1}(u_t) + c\left(\frac{t}{T}\right) \sigma_{u,t}^2 + e_t \\
&= a + \left[\sum_{i=1}^{n} \theta_i^b \, \varphi_i^b\left(\frac{t}{T}\right)\right] E_{t-1}(u_t) + \left[\sum_{i=1}^{n} \theta_i^c \, \varphi_i^c\left(\frac{t}{T}\right)\right] \sigma_{u,t}^2 + e_{nt} \\
&= a + \chi_{nt}^{b\prime} a_n^b + \chi_{nt}^{c\prime} a_n^c + e_{nt}, \qquad (7)
\end{aligned}$$

where $\chi_{nt}^b = \left[\varphi_1^b\left(\frac{t}{T}\right), \dots, \varphi_n^b\left(\frac{t}{T}\right)\right]' E_{t-1}(u_t)$, $\chi_{nt}^c = \left[\varphi_1^c\left(\frac{t}{T}\right), \dots, \varphi_n^c\left(\frac{t}{T}\right)\right]' \sigma_{u,t}^2$, $a_n^b = [\theta_1^b, \dots, \theta_n^b]'$, $a_n^c = [\theta_1^c, \dots, \theta_n^c]'$, $b_n\left(\frac{t}{T}\right) = \sum_{i=1}^{n} \theta_i^b \, \varphi_i^b\left(\frac{t}{T}\right)$, $c_n\left(\frac{t}{T}\right) = \sum_{i=1}^{n} \theta_i^c \, \varphi_i^c\left(\frac{t}{T}\right)$, and $e_{nt} = e_t + \left[a\left(\frac{t}{T}\right) - a\right] + \left[b\left(\frac{t}{T}\right) - b_n\left(\frac{t}{T}\right)\right] E_{t-1}(u_t) + \left[c\left(\frac{t}{T}\right) - c_n\left(\frac{t}{T}\right)\right] \sigma_{u,t}^2$.

Once $\chi_{nt}^{b\prime}$ and $\chi_{nt}^{c\prime}$ are constructed, a_n^b and a_n^c can be estimated by the least squares approach. In this regard, Andrews (1991a) demonstrates desirable asymptotic results for the estimates of a_n^b and a_n^c. It is straightforward to recover $b_n\left(\frac{t}{T}\right)$ and $c_n\left(\frac{t}{T}\right)$ with the estimates of a_n^b and a_n^c. Even if π_t and u_t are nonstationary, Park and Hahn (1999) show

that the time-varying cointegration coefficients can be estimated in a similar manner. Note that possible time-variation of π_t^* will be reflected in the disturbance term of the regression.

Since we cannot expand $b\left(\frac{t}{T}\right)$ and $c\left(\frac{t}{T}\right)$ with an infinite number of series functions, it is critical to decide n (the number of series functions) in the empirical analysis for obtaining a good approximation of $b\left(\frac{t}{T}\right)$ and $c\left(\frac{t}{T}\right)$. Regarding this issue, we use the h-block cross-validation (CV) and the modified h-block CV as selection criteria for n, as suggested by Burman, Chow, and Nolan (1994) as well as Racine (1997).[13] For a given block size (h), the h-block CV criterion can be expressed as:

$$CV = T^{-1}\sum_{t=h}^{T-h}(\pi_t - \hat{a}(t,h) - \chi_{nt}^{b'}\hat{a}_n^b(t,h) - \chi_{nt}^{c'}\hat{a}_n^c(t,h))^2, \tag{8}$$

where $\hat{a}(t,h)$, $\hat{a}_n^b(t,h)$, and $\hat{a}_n^c(t,h)$ are estimators of the coefficients in equation (7), obtained by removing the t-th observation, and the h observations preceding and following the t-th observation in the dependent and independent variables of the regression. The modified h-block CV criterion motivated by cases where $\frac{n}{T}$ is not negligible can be written as follows:

$$\begin{aligned} MCV = {} & T^{-1}\sum_{t=h}^{T-h}(\pi_t - \hat{a}(t,h) - \chi_{nt}^{b'}\hat{a}_n^b(t,h) - \chi_{nt}^{c'}\hat{a}_n^c(t,h))^2 \\ & + T^{-2}\sum_{t=h}^{T-h}\sum_{i=1}^{T}(\pi_i - \hat{a}(t,h) - \chi_{ni}^{b'}\hat{a}_n^b(t,h) - \chi_{ni}^{c'}\hat{a}_n^c(t,h))^2 \\ & + T^{-1}\sum_{i=1}^{T}(\pi_i - \hat{a} - \chi_{ni}^{b'}\hat{a}_n^b - \chi_{ni}^{c'}\hat{a}_n^c)^2 \end{aligned} \tag{9}$$

The n that minimizes CV or MCV is selected.

As shown in Table 5, we test various forms of series functions. Among them, the CV criterion is minimized when 1, m, m^2, $\cos(m)$, $\sin(m)$, $\cos(2m)$, $\sin(2m)$, $\cos(3m)$, and $\sin(3m)$ are used. Contrarily, the MCV criterion is minimized when 1, m, m^2, $\cos(m)$, $\sin(m)$, $\cos(2m)$, and $\sin(2m)$ are used. Hence, we conduct the estimation with

[13] The block size, h, is set as the integer nearest to $\frac{T}{6}$, as suggested by Burman et al. (1994).

both Fourier flexible forms to approximate $b\left(\frac{t}{T}\right)$ and $c\left(\frac{t}{T}\right)$ in the empirical analysis.

TABLE 5 h-block and modified h-block cross-validation criteria

	h-block CV	Modified h-block CV
$1, m, m^2$	4723.0	10361
$1, m, m^2, m^3$	5142.0	12066
$1, m, m^2, m^3, m^4$	5394.0	13413
$1, m, m^2, \ldots, m^5$	5561.0	14522
$1, m, m^2, \cos(m), \sin(m)$	1627.5	7291.2
$1, m, m^2, \cos(m), \sin(m),$ $\cos(2m), \sin(2m)$	1291.2	7109.9
$1, m, m^2, \cos(m), \sin(m), \ldots$ $, \cos(3m), \sin(3m)$	1134.8	7262.9
$1, m, m^2, \cos(m), \sin(m), \ldots$ $, \cos(4m), \sin(4m)$	1513.5	7695.9

Notes: h-block cross-validation and modified h-block cross-validation are data-dependent criteria for selecting the optimal number of series functions. Statistics for h-block cross-validation and modified h-block cross-validation are derived from equations (8) and (9), respectively. The n that minimizes the above CV criteria is selected. See Burman et al. (1994) and Racine (1997) for further discussion.

The estimates of b_t and c_t based on the nonparametric regression with the selected n minimizing MCV are plotted in Figure 2. We replace $E_{t-1}(u_t)$ with $u_t + \varsigma_t$ under the rational expectation assumption and the realized variance of u_t is used for $\sigma_{u,t}^2$, as displayed in the second row of Table 3. Interesting points emerge from Figure 2. First, the estimate of b_t is never significant during the entire sample period. Second, c_t is significantly positive between the early 1970s and early 1980s and becomes insignificant from the early 1980s until

the mid-2000s. Then, c_t briefly switches to be negative in the mid-2000s and becomes significantly positive after the GFC. Third, but most importantly, the evolution of b_t and c_t provide hints about the Fed's preference parameters. An insignificant b_t and significantly positive c_t between the early 1970s and early 1980s imply that $\kappa_t = 1$ along with non-zero values for α, λ_t, and γ_t during this period. The result suggests that it is difficult to explain movements of rates of inflation and unemployment using the Barro–Gordon model for the Fed's preference. Instead, the prediction from the Cukierman–Gerlach-type model is consistent with insignificant b_t and significantly positive c_t between the early 1970s and early 1980s. Since both b_t and c_t are insignificant between the early 1980s and the mid-2000s, the Cukierman–Gerlach-type model has a problem explaining these movements of b_t and c_t. Insignificant b_t and c_t are possible with the case of $\lambda_t = 0$ or the case of $\gamma_t \to 0$ and $\kappa_t = 1$, implying that the Fed is thinking of the inflation targeting. While neither the Barro–Gordon model nor the Cukierman–Gerlach-type model is consistent with estimation results between the early 1980s and mid-2000s, the inflation targeting is compatible with the results displayed in Figure 2. Except for the brief period when c_t becomes negative during the mid-2000s,[14] it has been significantly positive after the GFC, while b_t continues to be insignificant. These results after the mid-2000s suggest that the Fed again started placing more weights on positive deviations in the rate of unemployment from the target compared to negative deviations around the outbreak of the GFC.

In summary, the nonparametric regression suggests that the Cukierman–Gerlach-type model can approximately describe the Fed's preference before Volcker's period and following

[14] A negative c_t is possible when $\gamma_t < 0$, which implies that negative unemployment deviations from the target incur a greater loss in the Fed's preference than positive deviations. In this case, the Fed intends to generate a deflation bias, not an inflation bias.

the recent crisis. However, for the period of 1980s and 1990s, the inflation targeting is consistent with the estimation results of the nonparametric regression instead of the Cukierman–Gerlach-type model.

(Figure 2 here)

Figure 3 presents time-varying coefficients estimated with an alternative number of series functions based on the CV criterion for estimating equation (6). The results are similar to those in Figure 2. While b_t is insignificant during the entire sample period, c_t is significantly positive between the early 1970s and early 1980s, becomes insignificant until 2003:Q4, and then becomes significantly negative between 2004:Q1 and 2009:Q4. However, c_t becomes positive again at the end of the sample while being marginally significant for a brief period. Again, the movements of b_t and c_t are consistent with the Cukierman–Gerlach-type model between the early 1970s and early 1980s, and further with the inflation-targeting model between the early 1980s and early 2000s. The evolution of those parameters in the mid-2000s suggests that the Fed was shortly concerned about overheating before the GFC. Then, the results imply that the Cukierman–Gerlach-type model may represent the Fed's preference better than alternative models for the period after the GFC even though it shows marginal significance briefly.[15]

(Figure 3 here)

Figures 2 and 3 state that a single model for the Fed's preferences is rarely compatible

[15] Further, an estimation with regressors constructed from the ARMA-GARCH model is conducted. The evolutions of b_t and c_t are similar to those depicted in figures 2 and 3, but the majority of coefficients are insignificant. The insignificant result may arise from the generated regressor problem or possible parameter instabilities in the ARMA-GARCH model. The results are available upon request.

with the dynamics of inflation and unemployment throughout the entire sample period. Alternatively, Figures 2 and 3 suggest that the Fed's preferences have varied from the Cukierman–Gerlach-type asymmetry model during the 1970s and early 1980s, the inflation-targeting model during the mid-1980s, 1990s, and early 2000s, and to the Cukierman–Gerlach-type asymmetry model again after the GFC. This interpretation, based on Figures 2 and 3, implies that the changing stance of the Fed plays a significant role in understanding the macroeconomic dynamics of the US instead of the exogenous time-varying volatilities.

Figure 4 shows a comparison of the fitted values of the inflation rate from the constant-coefficient regression with values from the nonparametric regressions. The nonparametric regression presents a much tighter relationship between the inflation rate and the unemployment rate. Adjusted $\bar{R}^2$s increased from 0.0344 in the constant-coefficient regression to 0.6893 (based on the MCV) and 0.7451 (based on the CV) in the nonparametric regressions. These findings suggest that allowing the time-varying feature of the Fed's preference improves the understanding of the dynamics of the inflation rate and unemployment rate in the US.

(Figure 4 here)

6. DISCUSSION AND CONCLUSION

The results of this study provide supplementary evidence for the literature investigating the role of the Fed in understanding the dynamics of the rates of inflation and unemployment in the US. According to the findings from the constant-coefficient regression, the Cukierman–Gerlach model can be selected to represent the Fed's preference for the entire duration of the sample. Considering this result, the Fed may not have triggered the decline in the inflation rate since the early 1980s. Instead, an exogenous fall in the volatility of the unemployment rate around the mid-1980s led the Fed to perceive the lower marginal benefit of unexpected inflation, which has subsequently resulted in the stabilization of the rate of inflation. This interpretation

has similarities with Sims and Zha (2006) and other subsequent studies emphasizing the role of the volatility of innovations rather than changes in the monetary policy rule to explain the evolution of the inflation rate and unemployment rate.

However, econometric evidence in this study implies that coefficients in the reduced-form regression derived from the first-order condition of the Fed's optimization problem are not constant, indicating that the preference parameters of the Fed have evolved during the sample period. Considering this possibility, we run a nonparametric regression, and the results suggest that the Cukierman–Gerlach model can explain the relationship between the rates of inflation and unemployment between the early 1970s and early 1980s. This result implies that between the early 1970s and early 1980s, the Fed was more concerned with positive deviations from the natural rate of unemployment than negative deviations. The results of the nonparametric regression also suggest that the dynamic behavior of the inflation and unemployment rate from the early 1980s to the early 2000s is compatible with the inflation-targeting preference of the Fed. The starting point of this period roughly coincides with the beginning of Volcker's term. The inflation-targeting preference is also consistent with the consensus that the philosophy of the Volcker–Greenspan era at the Fed is very similar to inflation targeting (Bernanke and Mishkin (1997)). Moreover, the direction and timing of the movement in the Fed's preference parameters implied by the nonparametric regression are approximately consistent with the changes in the monetary policy rule, indicating that it has been more stabilizing since Volcker's term than before (Clarida et al. (2000)). This study also shows shifts in the Fed's preference during the 2000s and 2010s. The results imply that the Fed was briefly concerned about overheating before the GFC, and again started putting more weights on positive deviations of the unemployment rate from the target than negative deviations as in the Cukierman–Gerlach model after the GFC. Hence, the evidence from the nonparametric regression suggests that changes in the Fed's preference approximately

correspond with important changes in the monetary policy rule in the US between 1960:Q1 and 2017:Q3.

ACKNOWLEDGEMENT

We are grateful to conference participants at the 2016 Korea International Economic Association's Annual Meeting, the 2017 Korea Allied Economic Association's Annual Meeting and the 2017 Korea Money & Finance Association Annual Meeting for helpful discussions and comments. All remaining errors are ours. Sookyung Park acknowledges that this work was supported by a Korea University Grant.

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APPENDIX 1: NONSTATIONARY INFLATION RATE AND UNEMPLOYMENT RATE

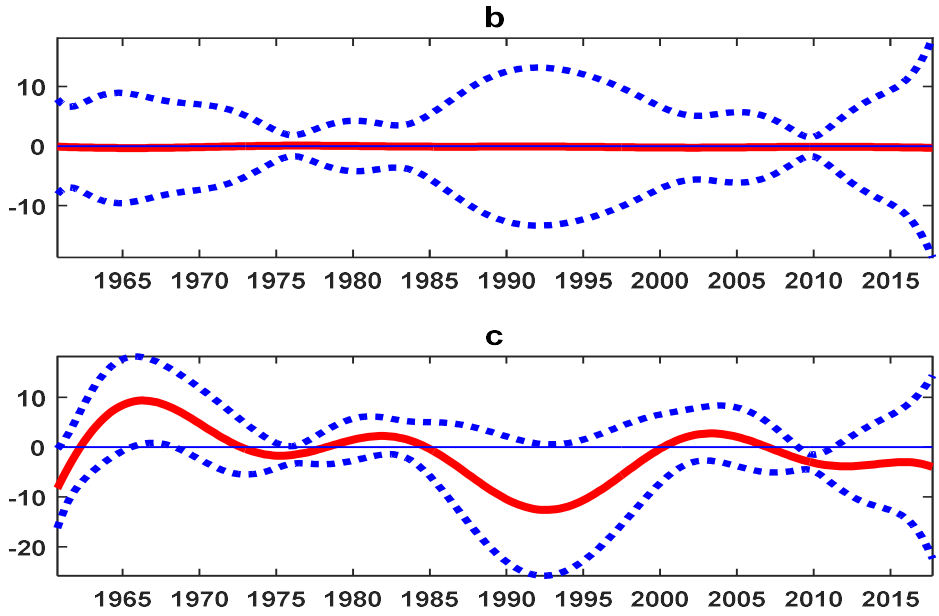
The results of the unit root tests in Table 2 show some possibility that the rates of inflation and unemployment are nonstationary. Hence, in this appendix, we assume that the inflation rate and unemployment rate are nonstationary and examine whether the two variables have a time-varying cointegration relationship. We conduct model specification tests proposed by Park and Hahn (1999), and Bierens and Martins (2010) to determine whether the two variables have a time-varying cointegration relationship instead of a constant one. As shown in Appendix Table 1, the model specification tests reject the constant cointegrating relationship in favor of the time-varying cointegration relationship. Based on this result, we plot the time-varying coefficients when unemployment follows the ARIMA (1,1,2) process. While b_t is insignificant during the entire sample period, c_t is significantly positive in the late 1960s, and then mostly insignificant up to 2017:Q3, except for a brief period in the late 2000s.

APPENDIX TABLE 1. Model specification tests for the time-varying cointegration

	τ_1	τ_2	Bierens and Martins test
$E_{t-1}(u_t) = u_t + \varsigma_t$ and $\sigma^2_{u,t}$ is the realized variance of u_t	187.5604	7.3468	28.1611 (0.0001)
$E_{t-1}(u_t)$ and $\sigma^2_{u,t}$ are from the ARIMA (1,1,2) model with GARCH (1,1) for u_t	163.0795	5.5312	27.1200 (0.0001)

Notes: Regarding τ_1, τ_2, and the Bierens and Martins (2010) test, the null hypothesis is a cointegration with constant coefficients, while the alternative hypothesis is a time-varying cointegration. The number of Chebyshev polynomials in the Bierens and Martins (2010) test and the number of time polynomials as superfluous regressors for τ_1 are set at three and five, respectively, however, the outcome is not sensitive to this choice. The numbers in parentheses in the Bierens and Martins (2010) test column are the p-values. The results in the third row of Appendix Table 1 are based on the assumption that $E_{t-1}(u_t)$ is formed from the ARIMA (1,1,2) specification for the rate of unemployment. The results are not sensitive to the choice of specification for the unemployment rate process. The 5% critical values for τ_1 and τ_2 reported in Park and Hahn (1999), and Shin (1994) are 11.07 and 0.22, respectively.

APPENDIX FIG. 1. Time-varying cointegrating coefficients



Notes: $\pi_t = a_t + b_t E_{t-1}(u_t) + c_t \sigma_{u,t}^2 + e_t$ is executed under the assumption that both inflation rate and unemployment rate are nonstationary and cointegrated. $E_{t-1}(u_t)$ and $\sigma_{u,t}^2$ are constructed from the ARIMA (1,1,2) model with GARCH (1,1) for u_t, as in Ruge-Murcia (2003).

FIGURE 1 Movements of inflation rate, unemployment rate, and the volatility of unemployment rate

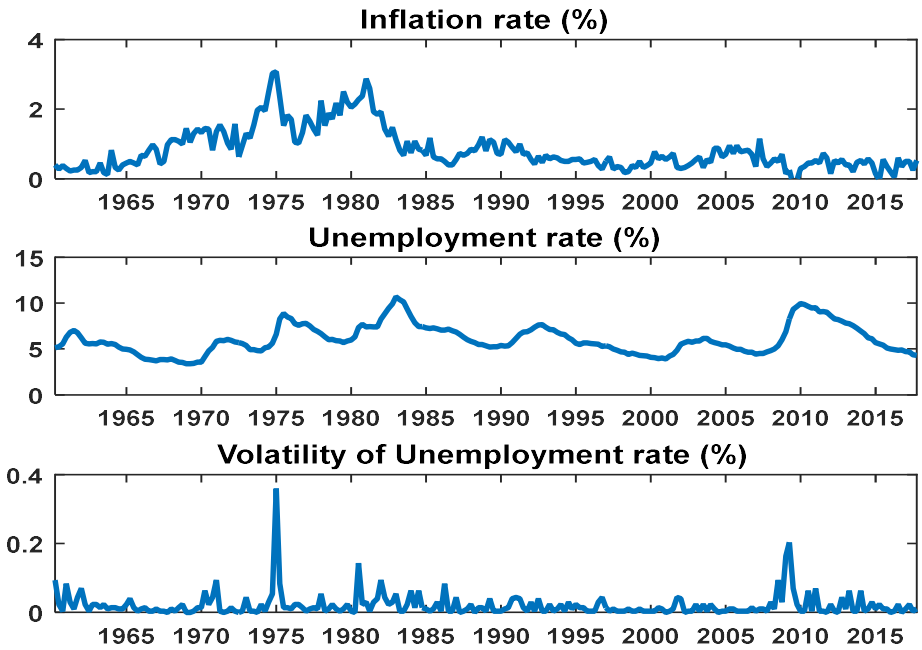
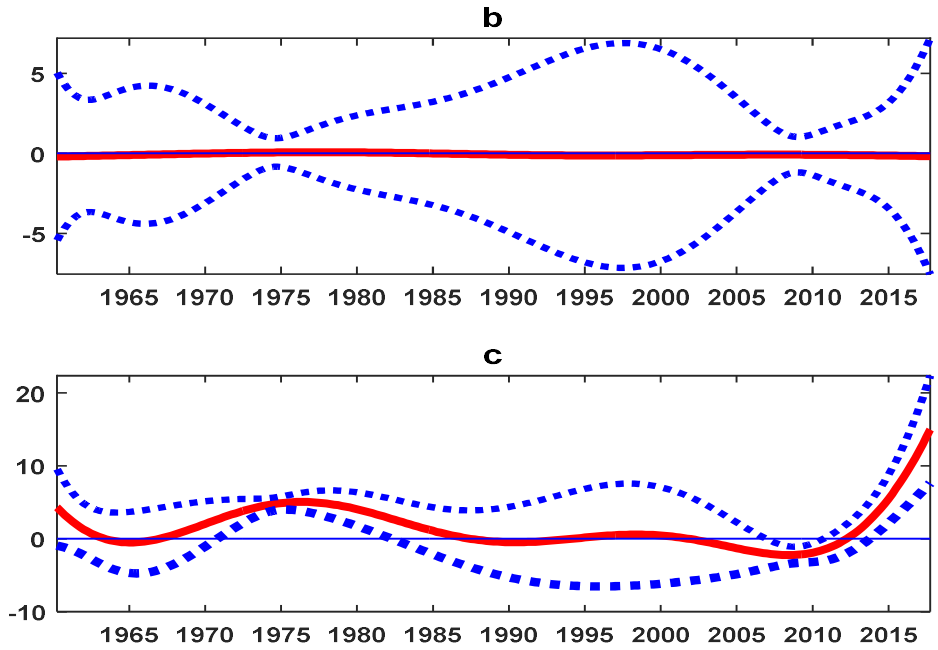
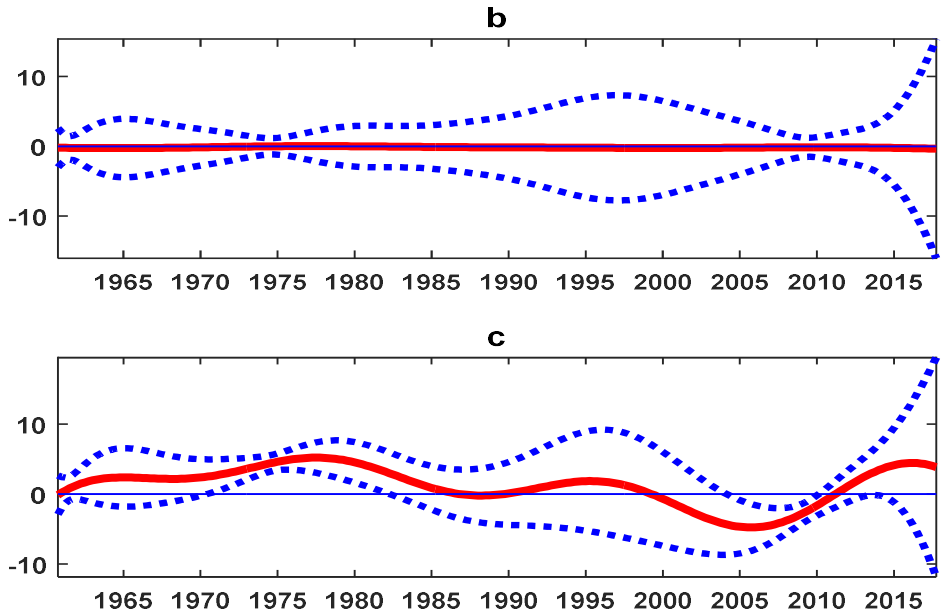


FIGURE 2 Time-varying coefficients from the nonparametric regression based on MCV criterion



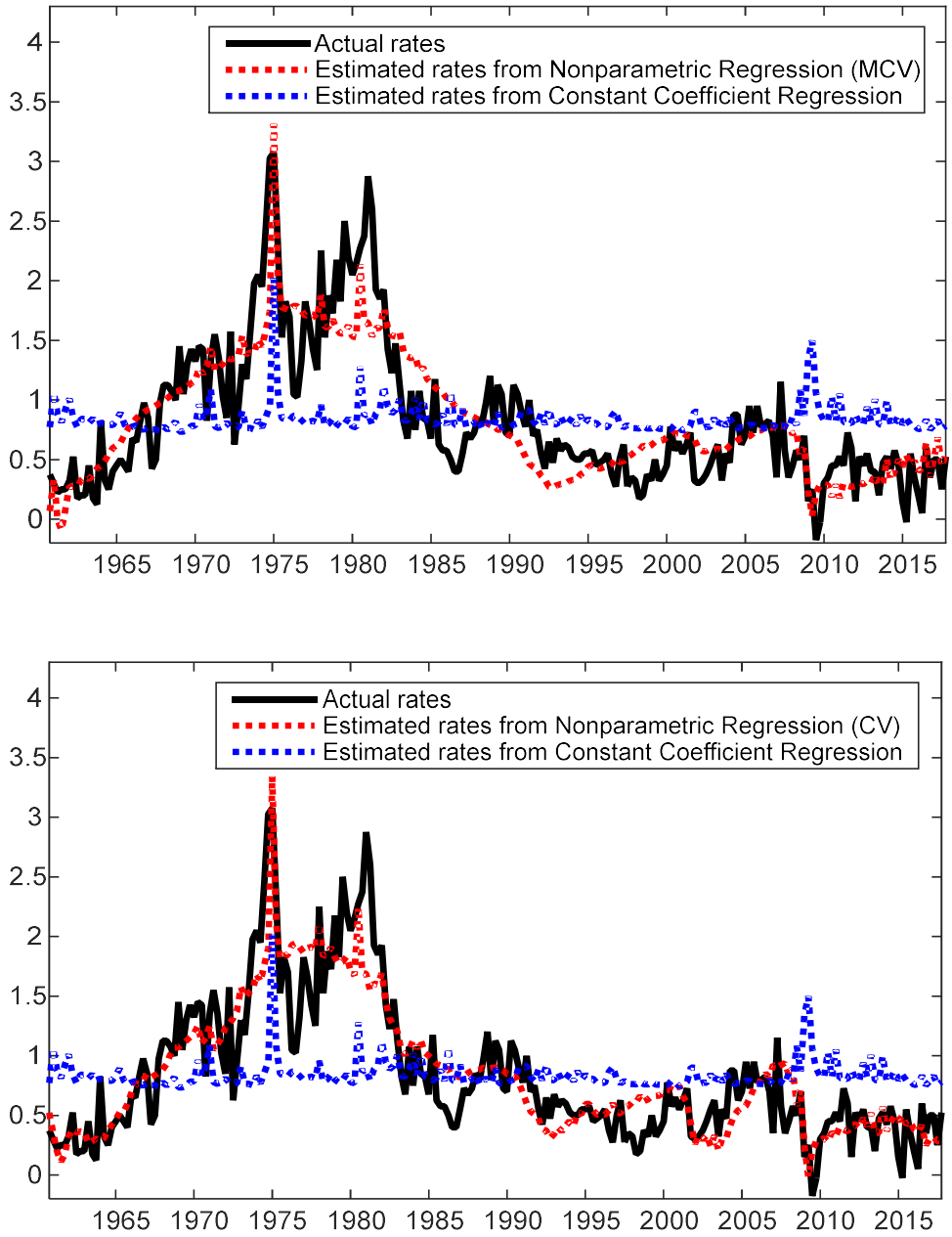
Notes: $\pi_t = a_t + b_t E_{t-1}(u_t) + c_t \sigma_{u,t}^2 + e_t$ is executed. $E_{t-1}(u_t)$ is replaced with $u_t + \varsigma_t$ (actual unemployment rate and forecast error) under the assumption of rational expectation, and the realized variance of u_t is used for $\sigma_{u,t}^2$. The number of series functions for estimating b_t and c_t is based on the modified h-block CV criterion shown in Table 5.

FIGURE 3 Time-varying coefficients from the nonparametric regression based on CV criterion



Notes: $\pi_t = a_t + b_t E_{t-1}(u_t) + c_t \sigma_{u,t}^2 + e_t$ is run. $E_{t-1}(u_t)$ is replaced with $u_t + \varsigma_t$ (actual rate of unemployment and forecast error) under the rational expectation assumption, and the realized variance of u_t is used for $\sigma_{u,t}^2$. As shown in Table 5, the number of series functions to estimate b_t and c_t is based on the h-block CV criterion.

FIGURE 4 Fitted values from constant coefficient regression and nonparametric regressions



Notes: The plots in each panel compare the actual rate of inflation with its fitted values from the constant coefficient regression and nonparametric regression, respectively. The fitted values of inflation rates from nonparametric regressions in the upper panel are based on the MCV criterion, while the CV criterion is used in the lower panel.