Trend Inflation and the Nature of Structural Breaks in the New Keynesian Phillips Curve

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Abstract

In this paper, we investigate the nature of structural breaks in inflation by estimating a version of the New Keynesian Phillips curve (NKPC) in the presence of a unit root in inflation. We show that, with a unit root in inflation, the NKPC implies an unobserved components model that consists of three components: a stochastic trend component, a component that depends upon current and future forecasts of real economic activity, and a stationary component which is potentially serially correlated (or a component of inflation that is not explained by the conventional forward-looking NKPC). Our empirical results suggest that, with an increase in trend inflation during the Great Inflation period, the response of inflation to real economic activity decreases and the persistence of the inflation gap increases due to an increase in the persistence of the unobserved stationary component. These results are in line with the predictions of Cogley and Sbordone (2008), who show that the coefficients of the NKPC are functions of time-varying trend inflation. Furthermore, the models presented in this paper display good in-sample and out-of-sample forecasting abilities relative to benchmark models in the literature that forecast inflation well.

Keywords: New Keynesian Phillips Curve, Trend Inflation, Inflation Gap, Unobserved Components Model, Structural Breaks.

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1. Introduction

While New Keynesian Phillips curves (NKPCs) are typically derived and estimated under the assumption that steady-state inflation or long-run trend inflation is close to zero, some authors have recently investigated the implications of non-zero or time-varying trend inflation on the dynamics of inflation within the New Keynesian framework. For example, Ascari (2004) shows that with non-zero trend inflation, both the long-run and the short-run properties of the New Keynesian models based on the Calvo pricing model can change dramatically. Cogley and Sbordone (2008) derive a version of the NKPC that incorporates a time-varying trend rate of inflation and show that the coefficients of the resulting NKPC are functions of the time-varying trend inflation. Ireland (2007) identifies the source of time-varying trend inflation as changes in the Federal Reserve Bank's implicit inflation target. Both Ireland (2007) and Cogley and Sbordone (2008) show that time-varying trend inflation contributes as a highly persistent component to post-war U.S inflation dynamics. Furthermore, Cogley et. al (2010) show that during the Great Inflation period of the 1970s, both the trend inflation and the persistence of the inflation gap increased considerably.

In the meantime, dating back to at least Nelson and Schwert (1977), there is substantial empirical evidence that U.S. inflation has a unit root. Based on reduced form unobserved components trend-cycle models of inflation with heteroscedastic shocks, Stock and Watson (2007) investigate the evolving nature of inflation dynamics. Harvey (2011) specifies a bivariate unobserved components model for inflation and real output based on a reduced form Phillips curve, with the lagged inflation terms replaced by a random walk. Lee and Nelson (2007) derive and estimate an unobserved components trend-cycle model for inflation and unemployment, as implied by the NKPC.

In this paper, we investigate the nature of structural breaks in inflation by estimating a version of the NKPC in the presence of a unit root in inflation. We first show that, with a unit root in inflation, the NKPC implies a reduced form unobserved components model of inflation that consists of three components: a stochastic trend component, a component that depends upon current and future forecasts of real economic activity, and a stationary component which is potentially serially correlated. Note that the last component can also be interpreted as the component of inflation that is not explained by the conventional forward-looking NKPC. We then investigate the nature and the implications of structural breaks in the latter two components, with a focus on the sensitivity of inflation to real economic activity and on the persistence of the component of inflation not explained by the conventional forward-looking NKPC. It is not our aim in this paper to directly investigate or identify the sources of structural breaks in these components. Rather, we ask whether the nature of structural breaks in these components is in line with the predictions of Cogley and Sbordone (2008), who derive a version of the time-varying NKPC in which the coefficients are the functions of time-varying trend inflation. Finally, the forecasting performances of the proposed models are evaluated.

Our empirical results can be summarized as follows. During the Great Inflation period of the 1970s, i) inflation was relatively less sensitive to real activity; ii) there was a surge in the persistence of the unobserved stationary component or the component not explained by the conventional forward-looking NKPC; and iii) there was an upward shift in trend inflation. All these results are consistent with the predictions of Cogley and Sbordone (2008). Based on observing an increased persistence in the unobserved stationary component alone, one cannot rule out the possibility that the backward-looking component or price indexation may have been relatively more important during the Great inflation period. However, when combined with the other two results mentioned above, this interpretation could be erroneous, as implied by Cogley and Sbordone (2008). We also find that our empirical models also forecast inflation well, suggesting that the relatively poor performance of recursive and rolling AR forecasts of inflation in the last decades may have been due to the disappearance in serial correlation of the unobserved stationary component in the Phillips curve in the early 1980s.

The outline of the paper is as follows. In Section 2, we provide model specifications. In particular, we first derive an unobserved components model of inflation as implied by the NKPC, and then explain how structural breaks can be incorporated into the model. In Section 3, we present empirical results. The forecasting performances of the proposed models are presented in Section 4. Section 5 provides concluding remarks.

2. Specification of an Empirical Model

Consider the following New Keynesian Phillips curve:

$$\pi_t = E_t(\pi_{t+1}) + kx_t + \eta_t, \tag{1}$$

where $E_t(.)$ refers to expectation formed conditional on information up to time t; x_t is a proxy for real economy activity such as the output gap, unemployment gap, or unit labor costs; and η_t is potentially serially correlated.

Serial correlation in the η_t term may reflect the importance of lagged inflation or higher order leads of expected inflation beyond $t+1$ in the NKPC. Although criticized for their lack of microeconomic foundations, lags of inflation are typically introduced by assuming some form of price indexation (Christiano et. al, 2005) or rule-of-thumb price-setting behavior (Gali and Gertler, 1999)[1]. Alternatively, Ascari (2004) and Cogley and Sbordone (2008) show that, in the presence of non-zero or time-varying trend inflation, the interaction between trend inflation and nonlinearities in the Calvo model of price adjustment results in additional leads of expected inflation in the NKPC. According to them, ignoring these multi-step expectation terms would lead to serial correlation in η_t, which may be mistakenly attributed to the lagged inflation terms.

We note that it is not our aim in this paper to directly investigate the source of serial correlation in the η_t term. Rather, we investigate the nature and the implications of structural breaks in the Phillips curve given above, with a focus on the structural breaks on the k parameter and the persistence of the η_t term. By iterating equation (1) in the forward direction, we have:

$$\pi_t = \lim_{j\to\infty} E_t(\pi_{t+j}) + k\sum_{j=0}^{\infty} E_t(x_{t+j}) + \tilde{z}_t, \tag{2}$$

[1] In the presence of price indexation or rule-of-thumb behavior, the hybrid NKPC is given by:

$$\pi_t = (1-\alpha)E_t(\pi_{t+1}) + kx_t + \alpha\pi_{t-1}, \ \ 0 \leq a \leq 1,$$

which can be rewritten as

$$\pi_t = E_t(\pi_{t+1}) + kx_t + \eta_t, \quad \eta_t = \alpha(\pi_{t-1} - E_t(\pi_{t+1})).$$

where $\tilde{z}_t = \sum_{j=0}^{\infty} E_t(\eta_{t+j})$. Here, the $\tilde{z}_t$ term is the component of inflation not explained by the conventional forward-looking NKPC. Note that in equation (2), the $\sum_{j=0}^{\infty} E_t(x_{t+j})$ term is a function of x_t. In the case that x_t is not exogenous, it is correlated with $\tilde{z}_t$, resulting in a difficulty in estimation of the model. For feasible estimation of the model, we replace $E_t(\pi_{t+j})$ with $E_{t-1}(\pi_{t+j})$ in equation (2) to get:

$$\pi_t = \lim_{j \to \infty} E_t(\pi_{t+j}) + k \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + z_t, \tag{3}$$

where $z_t = k(\sum_{j=0}^{\infty} E_t(x_{t+j}) - \sum_{j=0}^{\infty} E_{t-1}(x_{t+j})) + \tilde{z}_t$. Note that the first element of z_t is serially uncorrelated and it is a function of economic agents' revision on expectation of current and future real economic activities. Thus, potential serial correlation in z_t may be mainly due to serial correlation in the η_t term in (1).

In the presence of a unit root in inflation, the $\lim_{t \to \infty} E_t(\pi_{t+j})$ term, or the long-horizon inflation forecast, is the same as the random walk stochastic trend component of inflation. This is in the spirit of Beveridge and Nelson (1981). Depending upon whether we assume the output gap (x_t) to be observed or not, we have different model specifications. In what follows, we first consider a model specification in which the output gap is assumed to be observed. We then extend the model to the case of an unobserved output gap, which we extract from an unobserved components model of real output.

2.1 A Benchmark Model with Observed Output Gap: Model 1

In the specification below, we assume that the x_t and z_t terms can be approximated by finite-order autoregressive (AR) processes. For an AR(2) process for x_t and an AR(1) process for z_t, we have the following unobserved-components model which is consistent with the NKPC given in equation (1):

$$\pi_t = \pi_t^* + k \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + z_t, \tag{4}$$

$$\pi_t^* = \pi_{t-1}^* + v_t, \;\; v_t \sim i.i.dN(0, \sigma_v^2), \tag{5}$$

$$x_t = \phi_1 x_{t-1} + \phi_2 x_{t-2} + u_t,\ \ u_t \sim i.i.d.N(0, \sigma_u^2), \tag{6}$$

$$z_t = \psi z_{t-1} + \varepsilon_t,\ \ \varepsilon_t \sim i.i.d.N(0, \sigma_\varepsilon^2), \tag{7}$$

where u_t and ε_t are potentially correlated with correlation coefficient $\rho_{\varepsilon,u}$, as the z_t term involves a revision on expectation of the current and future real economic activities ($\sum_{j=0}^{\infty} E_t(x_{t+j}) - \sum_{j=0}^{\infty} E_{t-1}(x_{t+j})$). In the case that trend inflation is exogenous and is mainly driven by the long-run monetary policy of the central bank as in Cogley and Sbordone (2008), we can assume that the permanent shock to inflation is uncorrelated with both the transitory shock to inflation and the shock to output gap.

A state-space representation of the model is given by:

Measurement equation

$$\begin{bmatrix} \pi_t \\ x_t \end{bmatrix} = \begin{bmatrix} 1 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \pi_t^* \\ z_t \\ x_t^* \\ x_{t-1}^* \end{bmatrix} + \begin{bmatrix} k\sum_{j=0}^{\infty} E_{t-1}(x_{t+j}^*) \\ 0 \end{bmatrix} \tag{8}$$

Transition equation

$$\begin{bmatrix} \pi_t^* \\ z_t \\ x_t^* \\ x_{t-1}^* \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & \psi & 0 & 0 \\ 0 & 0 & \phi_1 & \phi_2 \\ 0 & 0 & 1 & 0 \end{bmatrix} \begin{bmatrix} \pi_{t-1}^* \\ z_{t-1} \\ x_{t-1}^* \\ x_{t-2}^* \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \\ 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} v_t \\ \varepsilon_t \\ u_t \end{bmatrix}, \tag{9}$$

$$\begin{bmatrix} v_t \\ \varepsilon_t \\ u_t \end{bmatrix} \sim i.i.d.N \left(\begin{bmatrix} 0 \\ 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_v^2 & 0 & 0 \\ 0 & \sigma_\varepsilon^2 & \rho_{\varepsilon,u}\sigma_\varepsilon\sigma_u \\ 0 & \rho_{\varepsilon,u}\sigma_\varepsilon\sigma_u & \sigma_u^2 \end{bmatrix} \right)$$

Note that the $\sum_{j=0}^{\infty} E_{t-1}(x_{t+j})$ term can be calculated by:

$$\sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) = F(I_2 - F)^{-1}\tilde{x}_{t-1}, \tag{10}$$

where $F = \begin{bmatrix} \phi_1 & \phi_2 \\ 1 & 0 \end{bmatrix}$ and $\tilde{x}_{t-1} = \begin{bmatrix} x_{t-1} \\ x_{t-2} \end{bmatrix}$.

2.1 A Benchmark Model with Unobbserved Output Gap: Model 2

When the output gap x_t is assumed to be unobserved, we can extract it from an unobserved components model of real output. Combined with the inflation equations, we have the following unobserved components model for inflation consistent with the NKPC:

$$\pi_t = \pi_t^* + k\sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + z_t, \tag{11}$$

$$\pi_t^* = \pi_{t-1}^* + v_t, \;\; v_t \sim i.i.d.N(0, \sigma_v^2), \tag{12}$$

$$z_t = \psi z_{t-1} + \varepsilon_t, \;\; \varepsilon_t \sim i.i.d.N(0, \sigma_\varepsilon^2). \tag{13}$$

$$y_t = \tau_t + x_t, \tag{14}$$

$$x_t = \phi_1 x_{t-1} + \phi_2 x_{t-2} + u_t, \;\; u_t \sim i.i.d.N(0, \sigma_u^2), \tag{15}$$

$$\tau_t = \mu + \tau_{t-1} + \eta_t, \;\; \eta_t \sim i.i.d.N(0, \sigma_\eta^2). \tag{16}$$

In the above specification, real output y_t is decomposed into a stochastic trend component τ_t that follows a random walk process with drift, and a cyclical component x_t that follows an AR(2) process. As before, all shocks in the model are assumed to be uncorrelated except for ε_t and u_t.

A state-space representation of the model is given by:

Measurement equation

$$\begin{bmatrix} \pi_t \\ y_t \end{bmatrix} = \begin{bmatrix} 1 & 1 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 1 \end{bmatrix} \begin{bmatrix} \pi_t^* \\ z_t \\ x_t \\ x_{t-1} \\ \tau_t \end{bmatrix} + \begin{bmatrix} k\sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) \\ 0 \end{bmatrix} \tag{17}$$

Transition equation

$$\begin{bmatrix} \pi_t^* \\ z_t \\ x_t \\ x_{t-1} \\ \tau_t \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \\ 0 \\ 0 \\ \mu \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & \psi & 0 & 0 & 0 \\ 0 & 0 & \phi_1 & \phi_2 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \pi_{t-1}^* \\ z_{t-1} \\ x_{t-1} \\ x_{t-2} \\ \tau_{t-1} \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} v_t \\ \varepsilon_t \\ u_t \\ \eta_t \end{bmatrix}, \tag{18}$$

$$\begin{bmatrix} v_t \\ \varepsilon_t \\ u_t \\ \eta_t \end{bmatrix} \sim i.i.d.N\left(\begin{bmatrix} 0 \\ 0 \\ 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_v^2 & 0 & 0 & 0 \\ 0 & \sigma_\varepsilon^2 & \rho_{\varepsilon,u}\sigma_\varepsilon\sigma_u & 0 \\ 0 & \rho_{\varepsilon,u}\sigma_\varepsilon\sigma_u & \sigma_u^2 & 0 \\ 0 & 0 & 0 & \sigma_\eta \end{bmatrix} \right)$$

As before, $\sum_{j=0}^{\infty} E_{t-1}(x_{t+j})$ is computed as:

$$\sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) = F(I_2 - F)^{-1}\tilde{x}_{t-1}, \tag{19}$$

where $F = \begin{bmatrix} \phi_1 & \phi_2 \\ 1 & 0 \end{bmatrix}$ and $\tilde{x}_{t-1} = \begin{bmatrix} x_{t-1} \\ x_{t-2} \end{bmatrix}$.

Incorporating Structural Breaks

We pretend to know the date of structural breaks in the output dynamics. That is, following Kim and Nelson (1999) and McConnell and Perez-Quiros (2000), we take the Great Moderation that started in 1984Q3 as a stylized fact. In addition, we allow for a productivity slowdown in 1973Q1. The structural breaks in the output dynamics are specified as follows:

(i) Productivity slowdown in 1973Q1 for real output: structural break in μ for Model 2 with unobserved output gap

$$\mu_{D_{1t}} = (1-D_{1t})\mu_0 + D_{1t}\mu_1, \;\; \mu_1 < \mu_0, \;\; D_{1t} = \{0,1\} \tag{20}$$

$$D_{1t} = \begin{cases} 0, & \text{if } t \leq 1973Q1, \\ 1, & \text{otherwise} \end{cases}$$

(ii) Great Moderation in 1984Q3 for real output: reduction in the variances of the shocks to real output

$$\sigma^2_{\eta,D_{2t}} = (1-D_{2t})\sigma^2_{\eta,0} + D_{2t}\sigma^2_{\eta,1}, \;\; \sigma^2_{\eta,1} < \sigma^2_{\eta,0}, \;\; D_{2t} = \{0,1\} \tag{21}$$

$$\sigma^2_{u,D_{2t}} = (1-D_{2t})\sigma^2_{u,0} + D_{2t}\sigma^2_{u,1}, \;\; \sigma^2_{u,1} < \sigma^2_{u,0}, \;\; D_{2t} = \{0,1\} \tag{22}$$

where

$$D_{2t} = \begin{cases} 0, & \text{if } t \leq 1984Q3, \\ 1, & \text{otherwise} \end{cases}$$

However, we do not pretend we know the dates of structural breaks in the dynamics of inflation. Our focus is on the nature of structural breaks in the sensitivity of inflation to current and future real economic activity and in the persistence and the variance of the component not explained by the forward-looking NKPC. Assuming that trend inflation is mainly driven by the long-run monetary policy of the central bank, one can conjecture that the evolution of trend inflation may be smooth and that there may be no abrupt shifts in the variance of the shocks to trend inflation. However, in order to test whether this is the case, we also consider structural breaks in the variance of the shocks to trend inflation.

An appropriate model selection procedure leads us to incorporate two structural breaks with unknown break dates for the dynamics of inflation. We thus employ a three-state Markov-switching model with absorbing states, as in Chib (1998). The conjecture is that the three regimes captured by the model are: i) a regime before the Great Inflation started; ii) the Great Inflation regime; and iii) a regime after the Volcker disinflation. The inflation dynamics with structural breaks are given by:

$$\pi_t = \pi_t^* + k_{S_t} \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + z_t, \tag{23}$$

$$\pi_t^* = \pi_{t-1}^* + v_t, \;\; v_t|S_t \sim i.i.dN(0, \sigma^2_{v,S_t}), \tag{24}$$

$$z_t = \psi_{S_t} z_{t-1} + \varepsilon_t, \;\; \varepsilon_t|S_t \sim i.i.d.N(0, \sigma^2_{z,S_t}), \tag{25}$$

$$S_t = 1,2,3$$

where S_t is a first-order Markov-switching variable with the following matrix of transition probabilities:

$$P = \begin{bmatrix} p_{11} & 1-p_{11} & 0 \\ 0 & p_{22} & 1-p_{22} \\ 0 & 0 & 1 \end{bmatrix}, \tag{26}$$

where the $(i,j)-th$ element refers to $Pr[S_t = j|S_{t-1} = i]$.

By incorporating the structural breaks outlined above, the complete empirical models can be specified as:

<u>Model 1 with Structural Breaks</u>

$$\pi_t = \pi_t^* + k \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + z_t, \tag{27}$$

$$\pi_t^* = \pi_{t-1}^* + v_t, \ \ v_t|S_t \sim i.i.d.N(0, \sigma^2_{v,S_t}), \tag{28}$$

$$z_t = \psi_{S_t} z_{t-1} + \varepsilon_t, \ \ \varepsilon_t|S_t \sim i.i.d.N(0, \sigma^2_{\varepsilon,S_t}), \tag{29}$$

$$x_t = \phi_1 x_{t-1} + \phi_2 x_{t-2} + u_t, \ \ u_t|D_{2t} \sim i.i.d.N(0, \sigma^2_{u,D_{2t}}), \tag{30}$$

where $cov(\varepsilon_t, u_t|D_{2t}, S_t) = \rho_{\varepsilon,u}\sigma_{u,D_{2t}}\sigma_{\varepsilon,S_t}$. All the other covariance terms are assumed to be zero.

Model 2 with Structural Breaks

$$\pi_t = \pi_t^* + k \sum_{j=0}^{\infty} E_{t-1}(x_{t+j}) + z_t, \tag{31}$$

$$\pi_t^* = \pi_{t-1}^* + v_t, \ \ v_t|S_t \sim i.i.d.N(0, \sigma^2_{v,S_t}) \tag{32}$$

$$z_t = \psi_{S_t} z_{t-1} + \varepsilon_t, \ \ \varepsilon_t|S_t \sim i.i.d.N(0, \sigma^2_{\varepsilon,S_t}), \tag{33}$$

$$y_t = \tau_t + x_t, \tag{34}$$

$$x_t = \phi_1 x_{t-1} + \phi_2 x_{t-2} + u_t, \ \ u_t|D_{2t} \sim i.i.d.N(0, \sigma^2_{u,D_{2t}}), \tag{35}$$

$$\tau_t = \mu_{D_{1t}} + \tau_{t-1} + \eta_t, \ \ \eta_t|D_{2t} \sim i.i.d.N(0, \sigma^2_{\eta,D_{2t}}), \tag{36}$$

where $cov(\varepsilon_t, u_t|D_{2t}, S_t) = \rho_{\varepsilon,u}\sigma_{u,D_{2t}}\sigma_{\varepsilon,S_t}$. All the other covariance terms are assumed to be zero.

3. Empirical Results: The Nature of Structural Breaks in Inflation Dynamics

For the inflation series, we use the annualized log-difference of the quarterly seasonally adjusted GDP chain-type price index. We choose the CBO (Congressional Budget Office) output gap series as a proxy for real economy activity in Model 1 and we extract measures of the output gap from the annualized log RGDP series in Model 2. The sample covers the period of 1952Q1-2007Q3 where the beginning of the sample is chosen to avoid large swings in inflation resulting from the Korean war, and the end of the sample marks the quarter prior to the 2007 financial crisis.

All data is taken primarily from the Federal Reserve Economic Database (FRED).

In determining the number of states, we consider model selection based on the Akaike Information Criterion (AIC). Then, we verify our model selection results by checking whether a given model captures most of the serial correlation and heteroskedasticity in the data. This is done by testing the null of no serial correlation in the standardized residuals and their squares. The number of structural breaks thus chosen is two[2].

Tables 1 and 2 contain the estimated parameters for Models 1 and 2 respectively. For each table, the second column reports the results for a model with structural breaks in the variances of the shocks to both the stochastic trend and stationary components. The third column reports the results for the model in which the shocks to the stochastic trend component is constrained to follow an i.i.d. process. Based on log-likelihood ratio tests, we cannot reject the null hypothesis that the variance of the permanent shocks is constant for each measure of inflation for both models. Thus, all of the discussion that follows is based on estimates of models with constant variance of permanent shocks. In addition, estimates of the unobserved components as well as the estimated regime probabilities are similar for Model 1 and 2, thus due to space considerations, only the estimates from Model 1 are shown and discussed. The empirical results can be summarized as follows.

First, the estimated transition probabilities imply that the two break dates are the early 1970s and the early 1980s. Figure 1 illustrates the estimated regime probabilities. As expected, the second regime as shown in both of these Figures roughly coincide with the Great Inflation period. The third regime roughly coincides with the period after the Volcker disinflation.

Second, the empirical models imply an upward shift in trend inflation during the Great inflation period. In Figure 2, the estimates of trend inflation and the 90% confidence intervals are depicted against the GDP Deflator. The ups and downs of trend inflation generally coincide with those of actual inflation, even though trend inflation is much smoother than actual inflation. It can be observed that trend inflation is low and steady in the early 1960s, started to rise in the late 1960s,

[2]For the chosen model, we could not reject the null hypothesis that the standardized residuals and their squares are white noise except for a short disinflationary period in the early 1980s.

peaking around 6% in the 1970s , fell after the Volcker disinflation of the early 1980s, and remained low and stable around 2% since the early 1990s. This general trajectory of trend inflation is similar to those reported elsewhere (e.g., Cogley and Sargent, 2002; Ireland, 2007; Cogley and Sbordone, 2008).

Third, estimates of the k parameters over the different regimes suggest that inflation was less sensitive to real economic activity in the 1970s than in the other regimes. The k parameter is statistically insignificant during the Great Inflation regime, while it is positive and statistically significant in the other two regimes. These results are consistent with those reported in Cogley et. al (2010) and Cogley and Sbordone (2008). The former report a decline in the coefficient on real economic activity during the 1970s based on recursive estimation of backward-looking Phillips curves. The latter report the same results based on estimation of a structural model with time-varying trend inflation.

Fourth, for both models, the persistence (ψ) and the variance of the unobserved stationary component (z_t) are much higher in the second regime (the Great Inflation) than in the other two regimes. Figure 3 depict the estimates of the inflation cycle and the inflation gap, where the inflation cycle is defined as the portion of inflation that is driven by current and future economic activity and the inflation gap as the sum of the inflation cycle and the unobserved stationary component. Note that the inflation gap diverges considerably from the inflation cycle during the Great Inflation period, as the result of an increase in the persistence and variance of the unobserved stationary component. Cogley and Sbordone (2008) and Cogley et. al (2010) also present evidence that inflation-gap persistence increased considerably during the Great Inflation and that it fell after the Volcker disinflation. However, these results are inconsistent with Stock and Watson (2007), who suggest that higher inflation persistence before the Volcker disinflation can be explained by a higher fraction of the variance of the change in inflation ($\Delta\pi_t$) explained by shocks to the stochastic trend component. As in Cogley and Sbordone (2008), we conjecture that the increase in the persistence and the variance of the unobserved stationary component (i.e. the portion of inflation not explained by the conventional forward-looking NKPC) during the Great Inflation period may reflect a failure to explicitly include additional leads of inflation expectation, the relative weight of which increased with an increase in trend inflation.

As for the estimated output dynamics, the sum of the AR coefficients associated with both the CBO output gap and the unobserved output gap as extracted from Model 2 suggest a highly persistent output gap. Note that the latter output gap series is consistent with the NKPC with time-varying trend inflation. In addition, both output gap series are quite volatile, and the standard deviation of shocks to both output gap series are roughly halved after the Great Moderation. According to Model 2, US trend output is deterministic across the postwar period [3]. This finding is consistent with Perron and Wada (2009), who show that once a one-time break in the slope function of U.S. trend output in 1973Q1 is allowed for, the trend function for equilibrium output is deterministic[4]. Finally, Model 2 implies an annual growth rate of U.S. real output of 3.82% before the productivity slowdown in 1973Q1, and 2.98% thereafter, which is roughly in line with other estimates in the literature.

A Comparison of Output Gaps

The output gap measures that enter the two unobserved components models are obtained from quite disparate methods. The CBO output gap in Model 1 is estimated from a large-scale multi-sector growth model, whereas the unobserved output gap in Model 2 is a byproduct of estimating a bivariate unobserved components model in inflation and real output that is consistent with the NKPC. Furthermore, the CBO output gap is a two-sided estimate, as it is constantly revised after new information about the macroeconomy becomes available in the data. On the other hand, estimates of the output gap obtained from Model 2 is a one-sided measure that only relies on (revised) data up until date t. Yet, from the empirical results reported in Tables 1 and 2, the sum of the

[3]The empirical model allows for a structural break in the variance of shocks to trend output due to the Great Moderation. However, we find that the variance of shocks to trend output is statistically insignificant both before and after the Great Moderation. Thus, the estimation results shown in Table 2 are based on a restricted model with constant trend output variance.

[4]To check the robustness of our results, we estimate Model 2 with two alternative specifications for the output trend function. However, we find that whether the slope term μ is specified as a constant or a random walk drift process, U.S. trend output is found to be deterministic.

AR coefficients and the standard deviation estimates of the shocks to both output gap series are remarkably similar.

Just how similar are these two output gap measures? Figure 4 presents a plot of the two output gaps series. Upon a quick glance, both output gaps contain movements that are generally similar. To better quantify the differences between the two, in Figure 5 we plot the CBO output gap alongside the 95% confidence interval associated with Model 2's estimates of the output gap. As shown, the CBO output gap is well contained within the 95% confidence set. In other words, the differences between the two output gap series are not statistically significant at the 5% significance level. Furthermore, we observe that the output gap associated with the bivariate unobserved components model has relatively narrow confidence bands when compared to other measures in the literature. These findings strengthens Kuttner (1994)'s case of using the bivariate unobserved components modeling approach as an effective shortcut method to obtain estimates of the output gap in place of a comprehensive supply-side analysis.

4. Inflation Forecasting

The forecasting performance of a model is often viewed as a useful metric for evaluating its empirical relevance. In this section, we conduct in-sample and out-of-sample inflation forecasting exercises with our proposed models to evaluate their ability to explain the data, as well as to evaluate the role of the output gap in producing inflation forecasts. We focus on forecasting $h-$period average inflation (at an annual rate), defined as $\pi_t^h = h^{-1}\sum_{i=0}^{h-1}\pi_{t-i}$. We use the notation $\pi_{t+h|t}^h$ to denote $h-$period-ahead forecasts of π_t^h made using data through time t.

We compare the forecasting performances of Models 1 and 2 with structural breaks against two univariate inflation forecasting models as described below.

Model 3: Atkeson and Ohanian (2001) inflation forecasting model:

$$\pi_{t+4|t}^4 = \pi_t^4 = \frac{1}{4}(\pi_t + ... + \pi_{t-3}). \tag{37}$$

The specification above, often referred to as the AO model, is well known in the literature for its ability to forecast inflation. It is a simple model that predicts the average four-quarter rate of inflation to be the same as the average rate of inflation over the previous four quarters[5]. Atkeson and Ohanian (2001) were the first to formally point out that since the mid- 1980s, four-quarter-ahead out-of-sample inflation forecasts obtained from such a specification have been more accurate than those implied by Phillips curve models. Based on more comprehensive analyses, Fisher et al. (2002), Stock and Watson (2003, 2007), and others confirm the AO result.

One important implication of the AO result is that since the mid- 1980s, information contained in measures of real economic activity has limited predictive content for inflation. Therefore, we also evaluate the forecasting performance of the following univariate unobserved components model, which is a restricted version of our proposed models with k set equal to zero.

Model 4: A univariate unobserved components model without the output gap ($k = 0$):

$$\pi_t = \pi_t^* + z_t, \tag{38}$$

$$\pi_t^* = \pi_{t-1}^* + v_t, \tag{39}$$

$$z_t = \psi_{S_t} z_{t-1} + \varepsilon_t, \tag{40}$$

$$\begin{bmatrix} v_t \\ \varepsilon_t \end{bmatrix} \sim i.i.d.N\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_v^2 & 0 \\ 0 & \sigma_{\varepsilon,S_t}^2 \end{bmatrix} \right). \tag{41}$$

Note that with $k = 0$ in Model 1 or Model 2, inflation is no longer influenced by movements in the output gap. Thus, by comparing the performance of the above specification to Models 1 and 2, we will be able to evaluate the role of the output gap in our inflation forecasts.

In our forecasting exercise, all four competing models are used to forecast three inflation series: the chain-weighted GDP deflator, CPI inflation, and the chain-weighted PCE inflation. We com-

[5] In Atkeson and Ohanian's (2001) paper, they define π_t^4 as the percentage change in the inflation rate between quarters $t-4$ and t. The specification here follows Stock and Watson's (2007) interpretation of the AO model.

pute both one-quarter-ahead in-sample inflation forecasts based on the full sample period 1952Q1-2007Q3, and four-quarter-ahead out-of-sample inflation forecasts for the sample 1994Q1-2007Q3 and the most recent period 2001Q1-2007Q3. For out-of-sample forecasting, a recursive procedure is followed. This means that for a forecast made at date t for date $t+4$, all estimation is made with data beginning in 1952Q1 through date t. Then, the procedure is repeated again with the same starting date but an expanding data window.

Table 3 reports the root mean squared errors (RMSEs) associated with the in-sample and out-of-sample inflation forecasts calculated from all four models. For the in-sample inflation forecasts, our proposed models 1 and 2, outperform the univariate models 3 and 4 for all inflation series. Similarly, for the out-of-sample case, our proposed models have lower RMSEs than the two competing univariate models.

To assess whether the differences in the out-of-sample predictive accuracies of our proposed models and the two univariate models are statistically significant, we analyze the out-of-sample inflation forecasting results using the modified Diebold-Mariano test statistic. The original Diebold-Mariano test statistic is a t-statistic associated with the null hypothesis that the mean squared errors of the two forecasts being compared is zero Diebold and Mariano (1995). The modified version as derived by Harvey et al. (1997) attempts to correct for the poor size property of the original test statistic in small samples.

The modified Diebold-Mariano test statistic are reported in Table 4 with their corresponding p-values in parentheses. From the evidence shown in the first three rows, neither our proposed models nor the competing univariate models have superior forecasting power for the out-of-sample period 1994Q1-2007Q3. Two exceptions are the RMSEs from our proposed models that are lower than the AO model at the 10% level of significance when forecasting the PCE and CPI inflation series. Focusing on the more recent period 2001Q1-2007Q3, however, our proposed models generally outperforms the competing univariate models.

In sum, the results presented in this section provide evidence that unobserved components models as implied by the NKPC can forecast inflation well, and are at times superior to univariate unobserved components models. This finding suggests that, in contrast to the generally accepted view in the literature, real economic activity still contains predictive content for inflation, and Phillips curve models remain useful for forecasting inflation.

5. Summary and Concluding Remarks

Ascari (2004) and Cogley and Sbordone (2008) show that non-zero or time-varying trend inflation results in additional leads of expected inflation in the NKPC. Cogley and Sbordone (2008) further show that the coefficients of the NKPC are functions of time-varying trend inflation. In this paper, we first derive a reduced-form model of inflation, as implied by the New Keynesian Phillips curve in the presence of a unit root in inflation. We then investigate the nature of structural breaks in the resulting unobserved components model, which consists of a stochastic trend component, a stationary component, and a component that depends upon current and future expected output gaps.

Our empirical results suggest that, with an increase in trend inflation during the Great Inflation period, the response of inflation to real economic activity decreases and the persistence of the inflation gap increases due to an increase in the persistence of the unobserved stationary component (i.e. the portion of inflation not explained by the conventional forward-looking NKPC). After the Volcker disinflation, the response of inflation to real economic activity increases and the persistence of the unobserved stationary component decreases. These results are all in line with the predictions of Cogley and Sbordone's (2008) model.

The increase in the persistence of the unobserved stationary component during the Great Inflation period may reflect a failure to explicitly include additional leads of inflation expectation, the relative weight of which might have increased with an increase in trend inflation, as suggested by Cogley and Sbordone (2008). One cannot rule out the possibility that the increased persistence of the inflation gap may be due to an increased importance of the backward-looking component. However, there exists no rigorous theory that may justify such time-varying importance of the backward-looking component. The empirical results in this paper could be interpreted as providing indirect evidence in support of Cogley and Sbordone (2008). They show that a purely forward-looking version of the NKPC fits the post-war U.S. data well after accounting for the effects of time-varying trend inflation on the multi-step expectations of inflation.

In explaining why inflation has become harder to forecast since the mid-1980s, Stock and Watson (2007) suggest that the changing auto-regressive (AR) coefficients and the deterioration of the low-order AR approximation accounts for the relatively poor performance of recursive and rolling AR forecasts. They show that this is because the variability of shocks to the permanent component of inflation has sharply declined since the mid 1980s. Our findings on the other hand suggest that the variability of shocks to trend inflation has remained unchanged. Instead, we find that the persistence as well as the variability of shocks to the unobserved stationary component has declined significantly since the mid 1980s, which may have been responsible for the relatively poor performance of recursive and rolling AR forecasts for inflation.

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References

Ascari, G., 2004. Staggered Prices and Trend Inflation: Some Nuisances. Review of Economics Dynamics 7, 642-667.

Atkeson, A., Ohanian, L.E., 2001. Are Phillips curves useful for forecasting inflation? Federal Reserve Bank of Minneapolis Quarterly Review 25(2), 2-11.

Beveridge, S., Nelson, C.R., 1981. A New Approach to the Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to Measurement of the 'Business Cycle'. Journal of Monetary Economics 7, 151-174.

Chib, S., 1995. Estimation and Comparison of Multiple Change-Point Models. Journal of Econometrics 86, 221-241.

Christiano, L., Eichenbaum, M., Evans, C., 2005. Nominal Rigidities and the Dynamic Effects of a shock to Monetary Policy. Journal of the Political Economy 113, 1-45.

Cogley, T., Sargent, T.J., 2002. Evolving Post-World War II U.S. Inflation Dynamics, in: Gertler M., Rogoff, K. (Eds.), NBER Macroeconomics Annual 2001. MIT Press: Cambridge.

Cogley, T., Primiceri, G.E., Sargent, T.J., 2010. Inflation-Gap Persistence in the US, American Economic Journal: Macroeconomics 2(1), 43-69.

Cogley, T., Sbordone, A.M., 2008. Trend Inflation, Indexation, and Inflation Persistence in the New Keynesian Phillips Curve. American Economic Review 98(5), 2101-2126.

Diebold, F.X., Mariano, R.S., 1995. Comparing predictive accuracy. Journal of Business & Economic Statistics 13(3), 253-263.

Fisher, J.D.M., Liu, C.T., Zhou, R., 2002. When can we forecast inflation? Federal Reserve Bank of Chicago Economic Perspectives 1Q/2002, 30-42.

Gali, J., Gertler, M., 1999. Inflation Dynamics: A Structural Econometric Analysis. Journal of Monetary Economics 44(2), 195-222.

Harvey, A.C., 2011. Modeling the Phillips Curve with Unobserved Components. Applied Financial Economics 21(1-2), 7-17.

Harvey, D.I., Leybourne, S.J. and Newbold, P., 1997. Testing the equality of prediction mean squared errors. International Journal of Forecasting 13, 281-291.

Ireland, P.N., 2007. Changes in the Federal Reserve's inflation target: causes and consequences. Journal of Money, Credit and Banking 39(8), 1851-2110.

Kim, C.-J., Nelson, C.R., 1999. Has the US Economy Become More Stable? A Bayesian Approach Based on a Markov-Switching Model of the Business Cycle. Review of Economics and Statistics, 81(4), 60816.

Kuttner, K.N., 1994. Estimating potential output as a latent variable. Journal of Business & Economic Statistics 12(3), 361-368.

Lee, J., Nelson, C.R., 2007. Expectation horizon and the Phillips curve: The solution to an empirical puzzle. Journal of Applied Econometrics 22, 161-178.

McConnell, M.M., Perez-Quiros, G.P., 2000. Output Fluctuations in the United States: What has Changed Since the Early 1980s? American Economic Review 90, 1464-1476.

Nelson, C.R., Schwert, G.W., 1977. Short-term Interest Rates as Predictors of Inflation: On Testing the Hypothesis that the Real Rate of Interest is Constant. American Economic Review 67(3), 478-486.

Perron, P., Wada, T., 2009. Let's take a break: Trends and cycles in US real GDP. Journal of Monetary Economics 56(6), 749-765.

Stock, J.H., Watson, M.W., 2003. Forecasting output and inflation: The role of asset prices. Journal of Economic Literature 41(3), 788-829.

Stock, J.H., Watson, M.W., 2007. Why Has U.S. Inflation Become Harder to Forecast? Journal of Money, Credit and Banking 39(1), 3-33.

Table 1: **Estimation results of Model 1 with observed output gap [1952Q1-2007Q3]**

Parameters	Estimates	
	Unconstrained Model	Constrained Model
	Output Gap Equation	
ϕ_1	1.168 (0.064)	1.166 (0.063)
ϕ_2	-0.240 (0.063)	-0.238 (0.062)
$\sigma_{u,0}$	1.034 (0.067)	1.034 (0.067)
$\sigma_{u,1}$	0.459 (0.034)	0.459 (0.034)
	Inflation Equation	
k_1	0.015 (0.008)	0.016 (0.008)
k_2	-0.014 (0.017)	-0.013 (0.017)
k_3	0.018 (0.010)	0.018 (0.010)
ψ_1	0.059 (0.215)	0.083 (0.191)
ψ_2	0.664 (0.134)	0.655 (0.136)
ψ_3	-0.072 (0.128)	-0.083 (0.130)
$\sigma_{\varepsilon,1}$	0.583 (0.088)	0.602 (0.077)
$\sigma_{\varepsilon,2}$	1.504 (0.181)	1.483 (0.172)
$\sigma_{\varepsilon,3}$	0.594 (0.061)	0.586 (0.060)
σ_v	-	0.357 (0.049)
$\sigma_{v,1}$	0.396 (0.089)	-
$\sigma_{v,2}$	0.295 (0.221)	-
$\sigma_{v,3}$	0.338 (0.061)	-
$\rho_{\varepsilon,u}$	-0.225 (0.082)	-0.225 (0.082)
p_{11}	0.988 (0.013)	0.988 (0.013)
$BreakDates$	1972Q1	1972Q1
p_{22}	0.979 (0.021)	0.979 (0.021)
$BreakDates$	1983Q3	1983Q3
$Log-likelihood$	-324.906	-325.081

Note: Standard Errors are in parentheses.

Table 2: **Estimation results of Model 2 with unobserved output gap [1952Q1-2007Q3]**

Parameters	Estimates	
	Unconstrained Model	Constrained Model
	Output Equation	
σ_η	0.000 (0.009)	0.000 (0.015)
μ_0	0.956 (0.041)	0.956 (0.042)
μ_1	0.744 (0.018)	0.744 (0.019)
ϕ_1	1.181 (0.062)	1.180 (0.065)
ϕ_2	-0.240 (0.063)	-0.238 (0.063)
$\sigma_{u,0}$	1.030 (0.067)	1.030 (0.067)
$\sigma_{u,1}$	0.466 (0.034)	0.466 (0.034)
	Inflation Equation	
k_1	0.015 (0.008)	0.016 (0.008)
k_2	-0.011 (0.013)	-0.011 (0.015)
k_3	0.015 (0.008)	0.016 (0.009)
ψ_1	0.049 (0.343)	0.078 (0.298)
ψ_2	0.677 (0.135)	0.668 (0.134)
ψ_3	-0.063 (0.104)	-0.076 (0.150)
$\sigma_{\varepsilon,1}$	0.579 (0.049)	0.600 (0.095)
$\sigma_{\varepsilon,2}$	1.501 (0.179)	1.479 (0.172)
$\sigma_{\varepsilon,3}$	0.597 (0.060)	0.588 (0.064)
σ_v	-	0.360 (0.053)
$\sigma_{v,1}$	0.402 (0.083)	-
$\sigma_{v,2}$	0.300 (0.218)	-
$\sigma_{v,3}$	0.339 (0.061)	-
$\rho_{\varepsilon,u}$	-0.232 (0.082)	-0.233 (0.082)
p_{11}	0.987 (0.013)	0.986 (0.014)
$BreakDates$	1970Q3	1970Q2
p_{22}	0.979 (0.021)	0.979 (0.021)
$BreakDates$	1982Q3	1982Q1
$Log-likelihood$	-327.938	-328.136

Note: Standard Errors are in parentheses.

Table 3: **RMSEs associated with in-sample and out-of-sample inflation forecasts**

	Forecasting Models			
Inflation Measure	Model 1	Model 2	Model 3	Model 4
One-step ahead in-sample forecasts, 1952Q1-2007Q3				
GDP Deflator	1.017	1.021	1.173	1.039
CPI	1.558	1.559	1.752	1.588
PCE	1.155	1.157	1.353	1.182
Four-step-ahead out-of-sample forecasts, 1994Q1-2007Q3				
GDP Deflator	0.477	0.476	0.488	0.491
CPI	0.858	0.891	0.922	0.910
PCE	0.665	0.669	0.690	0.679
Four-step-ahead out-of-sample forecasts, 2001Q1-2007Q3				
GDP Deflator	0.536	0.547	0.561	0.583
CPI	0.981	0.988	1.086	1.109
PCE	0.657	0.673	0.700	0.719

Note 1: Reported are the RMSEs from inflation forecasts. The RMSE statistic for the time period t_1 to t_2 is calculated as $RMSE_{t_1,t_2} = \sqrt{\frac{1}{T}\sum_{t=t_1}^{t_2}(\pi_{t+h}^h - \pi_{t+h|t}^h)^2}$, with $T = t_2 - t_1 - 1$. Note that $h = 1$ for in-sample one-quarter-ahead inflation forecasts, and $h = 4$ for out-of-sample four-quarter-ahead inflation forecasts.

Note 2: Model 1 is the unobserved components model with observed CBO output gap; Model 2 is the unobserved components model with unobserved output gap; Model 3 is Atkeson and Ohanian's (2001) inflation forecasting model; Model 4 is a univariate unobserved components model similar to Models 1 and 2 but with $k = 0$.

Table 4: **Evaluation of out-of-sample inflation forecasting performances**

	Competing Models			
Inflation Measure	Model 1 vs. Model 3	Model 2 vs. Model 3	Model 1 vs. Model 4	Model 2 vs. Model 4
Four-step-ahead out-of-sample forecasts, 1994Q1-2007Q3				
GDP Deflator	-0.387	-0.418	-0.276	-0.298
	(0.350)	(0.339)	(0.392)	(0.384)
CPI	-1.581	-0.603	-0.657	-0.236
	(0.059)	(0.274)	(0.257)	(0.407)
PCE	-1.314	-0.979	-0.223	-0.162
	(0.097)	(0.166)	(0.412)	(0.436)
Four-step-ahead out-of-sample forecasts, 2001Q1-2007Q3				
GDP Deflator	-0.583	-0.362	-1.487	-1.267
	(0.283)	(0.360)	(0.074)	(0.108)
CPI	-1.797	-1.339	-3.302	-2.357
	(0.042)	(0.096)	(0.001)	(0.013)
PCE	-2.072	-1.133	-1.732	-1.349
	(0.024)	(0.134)	(0.048)	(0.094)

Note 1: Reported are the modified Diebold-Mariano test statistics. In parentheses are their corresponding p-values under the null of equal predictive accuracy.

Note 2: Model 1 is the unobserved components model with observed CBO output gap; Model 2 is the unobserved components model with unobserved output gap; Model 3 is Atkeson and Ohanian's (2001) inflation forecasting model; Model 4 is a univariate unobserved components model similar to Models 1 and 2 but with $k = 0$.

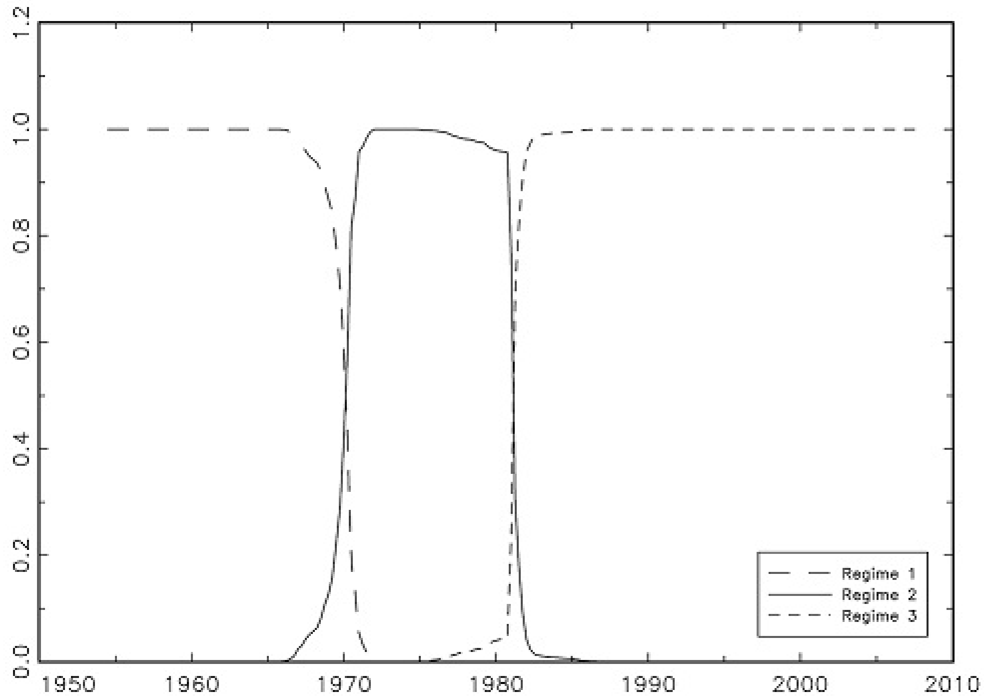

Figure 1: Smoothed Probabilities

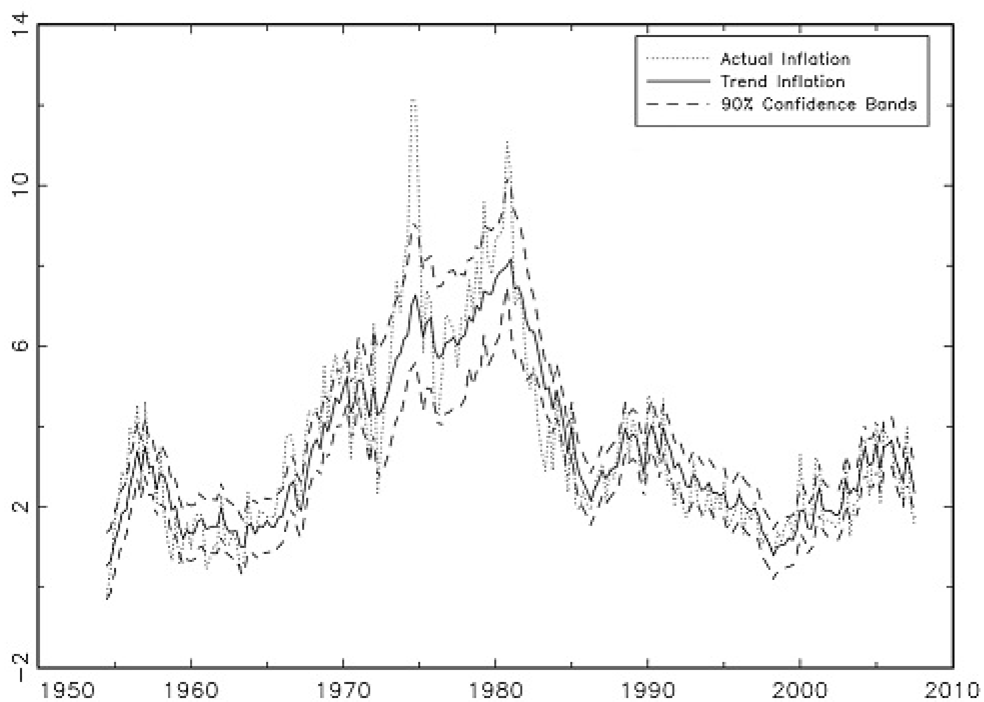

Figure 2: Actual Inflation and Trend Inflation

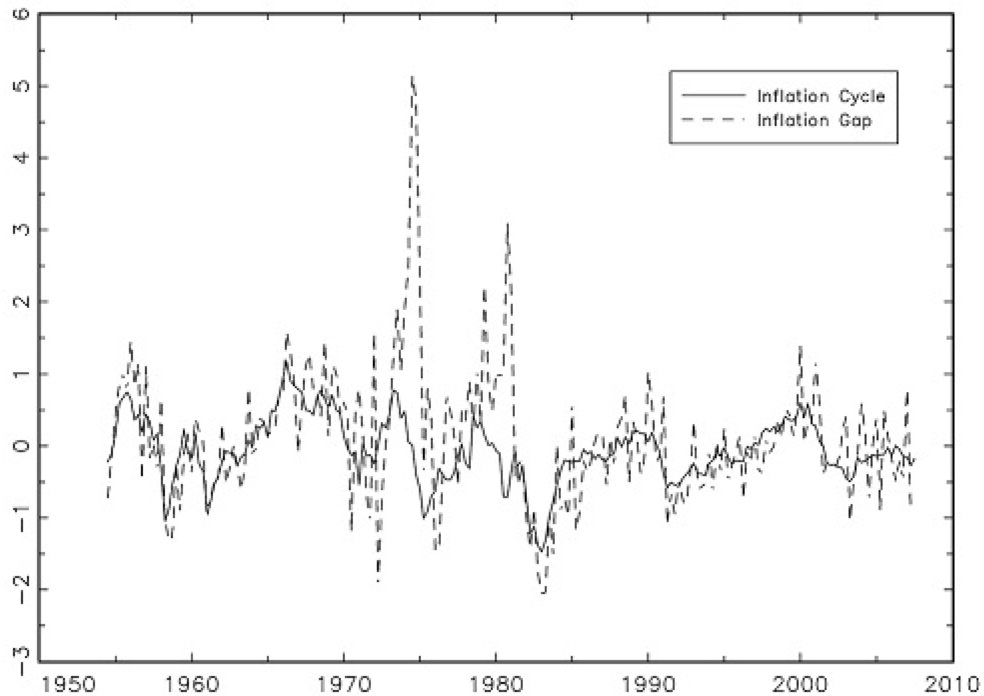

Figure 3: Inflation Cycle and Inflation Gap

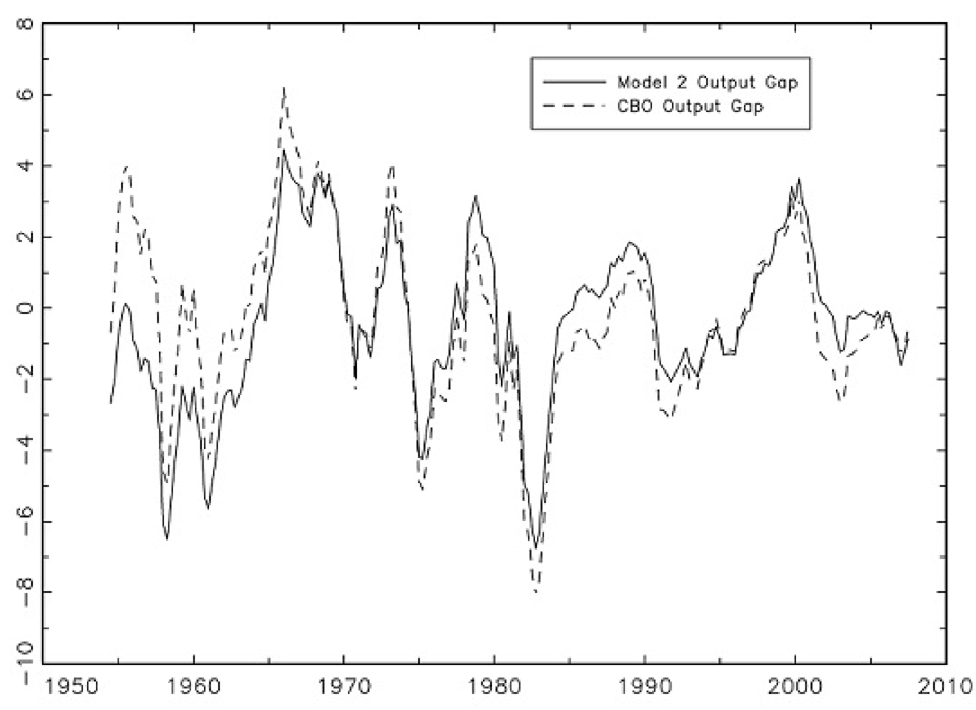

Figure 4: Model 2 Output Gap Estimates and the CBO Output Gap

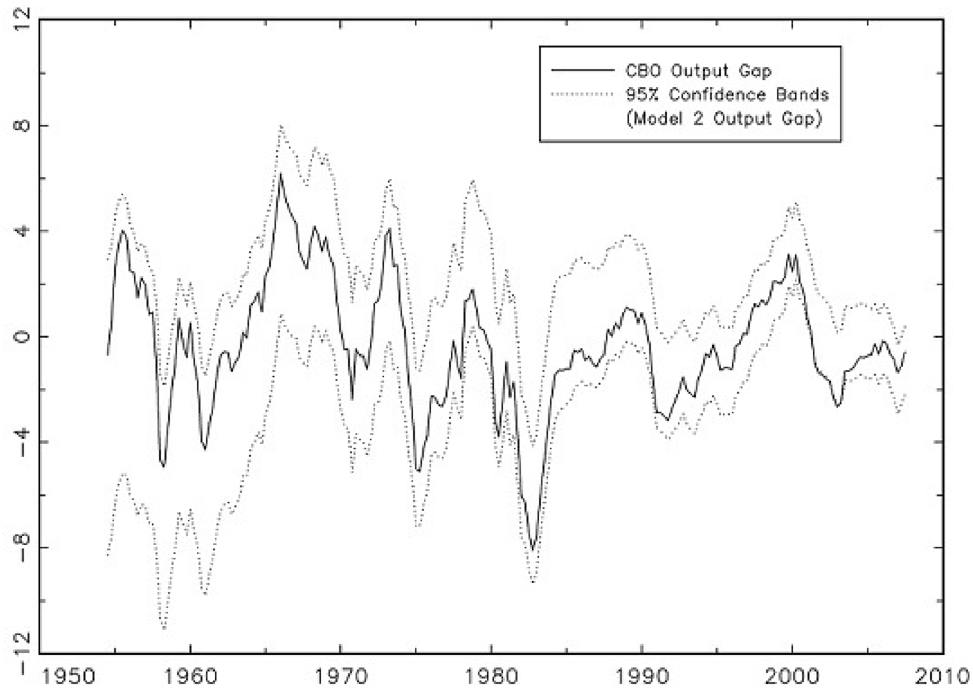

Figure 5: CBO Output Gap and Model 2 Output Gap's 95 Percent Confidence Bands