

Why has the U.S. economy stagnated since the Great Recession? *

Yunjong Eo[†]
University of Sydney

James Morley[‡]
University of Sydney

November 27, 2017

Abstract

Related to the idea of secular stagnation, the level and growth of U.S. real GDP have been much lower than would have been projected prior to the Great Recession. We investigate whether this stagnation is due to particularly large permanent effects of the Great Recession, a large and persistent negative output gap following the recession, or slower trend growth. Using a new Markov-switching time series model that allows a given recession and its recovery to be either U or L shaped and accounts for possible structural breaks in trend growth, we find the Great Recession was U shaped and the economy fully recovered from it by 2014. Instead, the relatively low level and growth of output appears to be driven by a reduction in trend growth that began in 2006 prior to the onset of the Great Recession. Our results help explain a lack of deflation in recent years without having to rely on a change in the slope of the Phillips Curve.

Keywords: Secular stagnation; The Great Recession; output gap; Markov switching; Phillips Curve

JEL classification: C22; C51; E32; E37

*The usual disclaimers apply.

[†]Yunjong Eo: School of Economics, University of Sydney, NSW 2006, Australia; Tel: +61 2 9351 3075; Email: yunjong.eo@sydney.edu.au

[‡]James Morley: School of Economics, University of Sydney, NSW 2006, Australia; Tel: +61 2 9351 3368; Email: james.morley@sydney.edu.au

1 Introduction

The slow growth of the U.S. economy in the wake of the Great Recession has led to a revival of earlier notions of secular stagnation ([Hansen, 1939](#)) and hysteresis ([Blanchard and Summers, 1986](#)). There are many different theories of secular stagnation, but [Summers \(2014, 2015\)](#) emphasizes the role of inadequate demand. Under his theory, the Global Financial Crisis (GFC) saw an unwinding of a financial bubble that had propped up the world economy. In its absence and in the face of the zero-lower-bound that prevents a further lowering of interest rates, inadequate demand causes the economy to grow at a slower rate than otherwise. This theory is related to the idea that inadequate demand resulting from the Great Recession may have produced hysteresis or even “super-hysteresis” effects ([Ball, 2014](#)) whereby a recession permanently lowers both the level and growth of economic activity. Using data from 23 countries, [Blanchard, Cerutti and Summers \(2015\)](#) document that many recessions have led to permanently lower output and growth, although they acknowledge that the causality could reflect supply shocks and financial crises producing both a recession and the subsequent stagnation. Meanwhile, [Cerra and Saxena \(2017\)](#) argue that all recessions have, on average, negative permanent effects on the level of output and question the relevance of the concept of an output gap in the first place, including its relevance for explaining weak economic activity and sluggish growth following the GFC.

A contrasting view of secular stagnation, emphasized by [Gordon \(2015\)](#), is that it reflects supply-side forces such as slower productivity growth and demographics that may have started before the Great Recession ([Fernald, 2015](#); [Fernald et al., 2016](#)). Notably, [Fernald et al. \(2017\)](#) use a growth accounting decomposition and find that, once allowing for cyclical effects, the slow growth in the U.S. economy since the Great Recession can be explained by slow growth of total factor productivity and the decline in labor force participation, with both phenomena starting before the onset of the recession and being unrelated to the financial crisis. Supporting this view, a number of empirical studies have documented a structural break in U.S. trend output growth in the mid 2000s prior to the Great Recession, including [Luo and Startz \(2014\)](#), [Grant and Chan \(2017\)](#), [Antolin-Diaz, Drechsel and Petrella \(2017\)](#), and [Kamber, Morley and Wong \(2017\)](#).

In this paper, we develop a highly flexible nonlinear time series model that allows us to examine the empirical support for competing views of why the level and growth of U.S. real GDP have been much lower than would have been projected prior to the Great Recession. In particular, we investigate whether this stagnation is due to particularly large permanent effects of the Great Recession, a large and persistent negative output gap following the recession, or slower trend growth. Building on [Hamilton \(1989\)](#), [Kim, Morley and Piger \(2005\)](#), and [Eo and Kim \(2016\)](#), our univariate Markov-switching model of output growth allows a given recession and its recovery to be either U or L shaped. We also consider possible structural breaks in trend growth. In particular, using the testing procedure in [Qu and Perron \(2007\)](#), our empirical analysis supports a reduction in trend growth in 2006:Q1. Allowing for this break in our Markov-switching model, we find that the Great Recession was U shaped and the economy fully recovered from it by 2014. Thus, the stagnation of U.S. real GDP appears to be driven primarily by a reduction in trend growth that began prior to the onset of the Great Recession. Consistent with the findings in [Huang and Luo \(2017\)](#), our results also help explain a lack of deflation in recent years without having to rely on a change in the slope of the Phillips Curve.

Our analysis is related to [Huang, Luo and Startz \(2016\)](#), who also consider a univariate time series model with two different types of recessions, but determine the prevailing regime using NBER dates and assume a given recession is predetermined as being either L or U shaped. Our Markov-switching model is more directly an extension of [Hamilton \(1989\)](#) and [Kim, Morley and Piger \(2005\)](#) to allowing two different types of recessions by modeling regimes as being stochastic. We believe this is a more natural assumption given that the exact timing and nature of future recessions is not very predictable in practice. This also leads to a different result than [Huang, Luo and Startz \(2016\)](#) in terms of categorizing the Great Recession as being U shaped rather than L shaped. Our model is also somewhat related to [Kim and Murray \(2002\)](#), [Kim and Piger \(2002\)](#), and [Kim, Piger and Startz \(2007\)](#), who consider multivariate unobserved components models with Markov-switching in both the trends and cycles of panels of macroeconomic time series, thus allowing recessions to be L or U shaped. However, our model is univariate and much more parsimonious. Also, by considering a growth rate specification, as in [Kim, Morley and Piger \(2005\)](#), instead of an

unobserved components structure, we do not impose a strong and potentially false restriction on the correlation between underlying permanent and transitory shocks (see [Morley, Nelson and Zivot \(2003\)](#)).

The rest of this paper proceeds as follows. In Section 2, we present the details of our new Markov-switching model and show how it can generate L and U shaped recessions. In Section 3, we present estimates for a benchmark specification of our model that allows a reduction in trend growth in 2006:Q1. As part of our analysis, we compare the realized path of U.S. real GDP to what would have been projected prior to the Great Recession. We also characterize all postwar recessions as being either L or U shaped. Then we estimate the output gap based on our model using the trend-cycle decomposition method for regime-switching models proposed in [Morley and Piger \(2008\)](#). In Section 4, we consider structural change in trend growth in more detail, testing for the existence of structural breaks and examining the influence of alternative assumptions about structural breaks on our inferences as to why the level and growth of U.S. real GDP have been much lower than would have been projected prior to the Great Recession. As part of this analysis, we consider the forecasting relationships between the estimated output gaps for different assumptions about structural breaks and inflation. Section 5 concludes.

2 A Markov-Switching Model with Two Types of Recessions

We consider a univariate nonlinear time series model of postwar U.S. real GDP that captures business cycle asymmetry. The model extends the Markov-switching models in [Hamilton \(1989\)](#) and [Kim, Morley and Piger \(2005\)](#) by allowing for two different types of contractionary regimes: (i) an L-shaped regime with permanent effects on the level of output, as in [Hamilton \(1989\)](#) and (ii) a U-shaped regime with only transitory effects, as in [Kim and Nelson \(1999a\)](#). This structure of allowing a given recession and its recovery to be either U or L shaped is motivated by [Eo and Kim \(2016\)](#), who find a Markov-switching model with time-varying regime-dependent mean growth rates that depend on each other across booms and recessions

fits the U.S. data better than the simpler Markov-switching models in [Hamilton \(1989\)](#) and [Kim, Morley and Piger \(2005\)](#). Building on the specification in [Kim, Morley and Piger \(2005\)](#), our extended model for real GDP growth, Δy_t , is given as follows:

$$\begin{aligned} \Delta y_t = \mu_0 &+ \mu_1 \mathbf{1}(S_t = 1) + \lambda_1 \sum_{k=1}^m \mathbf{1}(S_{t-k} = 1) \\ &+ \mu_2 \mathbf{1}(S_t = 2) + \lambda_2 \sum_{k=1}^m \mathbf{1}(S_{t-k} = 2) + e_t, \end{aligned} \quad (1)$$

where $e_t \sim i.i.dN(0, \sigma^2)$, S_t is a latent Markov-switching state variable that takes on discrete values of 0, 1, and 2 such that $S_t = 0$ for the expansionary regime, $S_t = 1$ for the L-shaped contractionary regime, $S_t = 2$ for the U-shaped contractionary regime according to transition probabilities $Pr[S_t = j | S_{t-1} = i]$ for $i, j = 1, 2, 3$, and $\mathbf{1}(\cdot)$ is an indicator function. Note that, following the results in [Hamilton \(1989\)](#), [Kim, Morley and Piger \(2005\)](#), [Morley and Piger \(2012\)](#), [Huang, Luo and Startz \(2016\)](#), [Eo and Kim \(2016\)](#), and others, we do not include linear autoregressive dynamics in what is already a reasonably highly-parameterized model.¹

The λ_1 and λ_2 parameters for the two contractionary regimes imply the possibility of ‘bounce-back effects’, as in [Kim, Morley and Piger \(2005\)](#). However, we assume that only an L-shaped regime can have permanent effects on the level of output, while a U-shaped regime does not.² In particular, to differentiate the two regimes, we impose restrictions on λ_2 and the transition probabilities.³

First, the bounce-effect parameter λ_2 for the U-shaped regime is constrained such that

¹Specifically, these earlier papers find that linear autoregressive dynamics are not particularly important once allowing for a Markov-switching mean. However, it is important to note that [Morley and Piger \(2012\)](#) find statistical support for Markov-switching nonlinearity based on the [Carrasco, Hu and Ploberger \(2014\)](#) test and a bootstrap likelihood ratio test given bounce-back specifications from [Kim, Morley and Piger \(2005\)](#) (though not for the [Hamilton \(1989\)](#) specification with only an L-shaped contractionary regime) when allowing for AR(2) dynamics in U.S. real GDP growth under the null of linearity.

²In addition to our consideration of a latent Markov-switching state variable instead of predetermined NBER dates, this assumption is the other key distinction from [Huang, Luo and Startz \(2016\)](#), who assume no bounce-back effect for their L-shaped regime, which is what we find in practice, but allow for possible permanent effects with their U-shaped regime.

³Typically with Markov-switching models, it is necessary to place a labelling restriction such as $\mu_1 < 0$ and $\mu_2 < 0$ to identify the model. However, because there is no bounce-back term when $S_{t-k} = 0$, the model is identified given the restriction on λ_2 and the transition probabilities only. Thus, we place no restrictions on the other conditional mean parameters in (1).

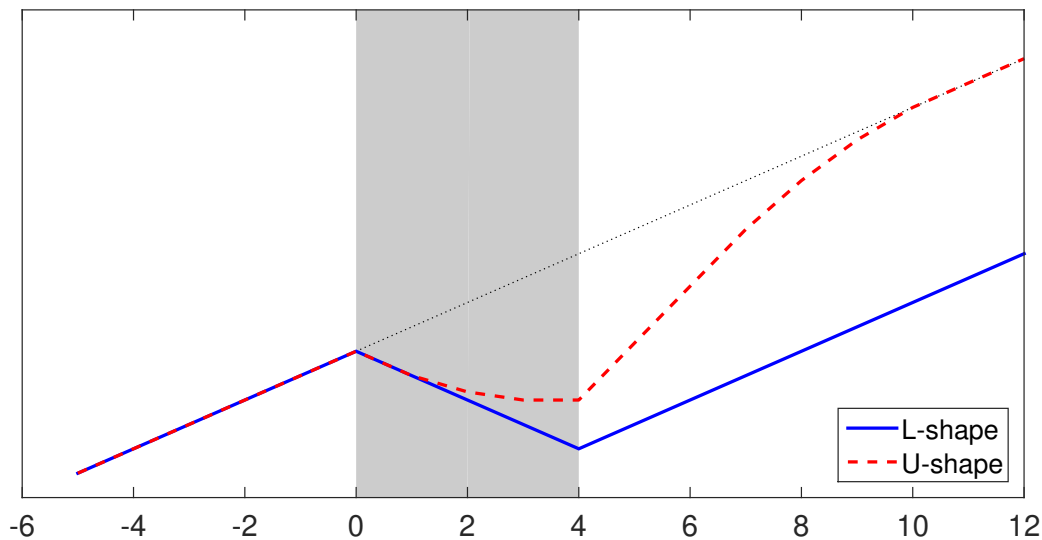
$\mu_2 + m \cdot \lambda_2 = 0$, so that the bounce-back effect $m \cdot \lambda_2$ exactly cancels out the contractionary effect μ_2 . By contrast, λ_1 for the L-shaped regime is left unrestricted. As long as $\mu_1 + m \cdot \lambda_1 < 0$, an L-shaped contraction has a permanent, negative effect on the level of output. Meanwhile, if $\lambda_1 = 0$, the L-shaped regime is equivalent to the original restricted version of a contractionary regime considered in [Hamilton \(1989\)](#) that does not allow any bounce-back effect.

Second, we assume that the economy does not switch between L-shaped and U-shaped regimes without going through an expansionary regime first. This sequencing of regimes can be imposed using restrictions on the regime transition probabilities as follows: $Pr[S_t = 2|S_{t-1} = 1] = 0$ for an L-shaped regime to U-shaped regime transition and $Pr[S_t = 1|S_{t-1} = 2] = 0$ for a U-shaped regime to L-shaped regime transition, $Pr[S_t = 1|S_{t-1} = 0] = q_1$ for an expansionary regime to L-shaped recession regime transition, and $Pr[S_t = 2|S_{t-1} = 0] = q_2$ for an expansionary regime to U-shaped regime transition. The contractionary regime continuation probabilities are specified as follows: $Pr[S_t = 1|S_{t-1} = 1] = p$ and $Pr[S_t = 2|S_{t-1} = 2] = r$. Thus, the regime transition matrix is given by

$$\Pi = \begin{bmatrix} 1 - q_1 - q_2 & 1 - p & 1 - r \\ q_1 & p & 0 \\ q_2 & 0 & r \end{bmatrix}. \quad (2)$$

Figure 1 illustrates how the two different types of contractionary regimes create different types of recessions in terms of their long-run effects on the level of output. For this demonstration, we simulate the model before, during, and after the occurrence of the two types of contractionary regimes. We set the length of the post-recession ‘bounce-back’ to $m = 6$ periods and the model parameters to be $\mu_0 = 1$ for the expansionary regime, $\mu_1 = -2$ and $\lambda_1 = 0$ for the L-shaped regime, and $\mu_2 = -2$ and $\lambda_2 = -\mu_2/m$ for the U-shaped regime. For clarity in seeing the impact of the relative impact of the two different regimes, we abstract from the e_t shocks in the model in (1) when simulating the path of output. In both cases, we assume that the economy is hit by a contractionary regime at time $t = 0$ that lasts for 4 quarters and causes a recession. For the U-shaped regime, the bounce-back term takes effect as the recession continues and flattens out the path of output, with the economy then

Figure 1: Illustration of L-shaped and U-shaped Recessions



Note: The shaded area denotes the contractionary regime.

growing quickly and recovering to its pre-recession path after the recession is over. In this sense, the recession has no permanent effect on the level of output and appears U shaped. By contrast, for the L-shaped regime, the absence of a bounce-back effect means that the economy contracts sharply in the recession and does not recover to its pre-recession path after the recession is over, but only grows at the usual expansionary rate. Thus, this recession has a permanent effect on the level of output and appears L shaped.

3 Estimates for a Benchmark Specification

The raw data are seasonally adjusted quarterly U.S. real GDP for the sample period of 1947:Q1 to 2017:Q2 and were obtained from the St. Louis Fed (FRED) database. Quarterly growth rates are calculated as 100 times the first differences of the natural logarithms of the level data. We estimate the model using output growth starting in 1947:Q2 and maximum likelihood estimation (MLE). The likelihood is constructed based on the filter presented in [Hamilton \(1989\)](#) and keeping track of 3^m states in each period. The length of the post-recession ‘bounce-back’ is set to $m = 6$ quarters following [Kim, Morley and Piger \(2005\)](#). Standard errors are based on numerical second derivatives.

Table 1: Maximum Likelihood Estimates for the Benchmark Specification

Parameter	Estimate	S.E.
q_1	0.0285	(0.0224)
q_2	0.0334	(0.0174)
p	0.7354	(0.1289)
r	0.8020	(0.0851)
σ^2	0.4370	(0.0500)
μ_0	0.9570	(0.0755)
μ_1	-1.1038	(0.4219)
λ_1	-0.0170	(0.0948)
μ_2	-1.9554	(0.1864)
δ	-0.5197	(0.1361)
log-lik	-342.47	

Notes: The benchmark specification allows for a structural break in trend growth in 2006:Q1 in (3). The standard errors of the parameter estimates are reported in parentheses. The estimation sample period is 1947:Q2 to 2017:Q2.

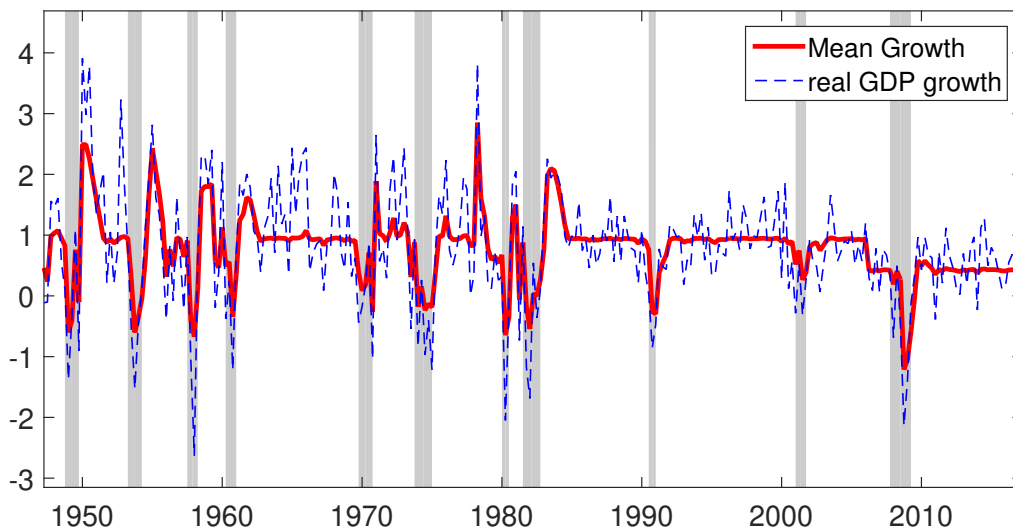
For our benchmark specification, we allow for a structural break in trend growth in 2006:Q1 following Luo and Startz (2014) and Fernald et al. (2017), but we examine different possibilities for the timing of a slowdown in more detail in Section 4. To account for the structural break, the model specification in (1) is altered accordingly:

$$\begin{aligned} \Delta y_t = \mu_0 + \delta \mathbf{1}(t > T_b) &+ \mu_1 \mathbf{1}(S_t = 1) + \lambda_1 \sum_{k=1}^m \mathbf{1}(S_{t-k} = 1) \\ &+ \mu_2 \mathbf{1}(S_t = 2) + \lambda_2 \sum_{k=1}^m \mathbf{1}(S_{t-k} = 2) + e_t, \end{aligned} \quad (3)$$

where the break date T_b for our benchmark specification is 2006:Q1 and δ is left unrestricted in estimation.

Table 1 presents maximum likelihood estimates for the independently-determined parameters of the benchmark specification. The estimated output growth rates $\hat{\mu}_0 + \hat{\mu}_1 < 0$ for the L-shaped regime and $\hat{\mu}_0 + \hat{\mu}_2 < 0$ for the U-shaped regime indicate that both regimes are clearly contractionary, although again this was not imposed in estimation. The estimated bounce-back parameter $\hat{\lambda}_1 = -0.0170$ for the L-shaped regime is very close to zero and insignificant. This implies that the L-shaped regime for the benchmark specification is essentially

Figure 2: Quarterly Output Growth and Time-Varying Mean



Note: The shaded areas denote NBER recession dates.

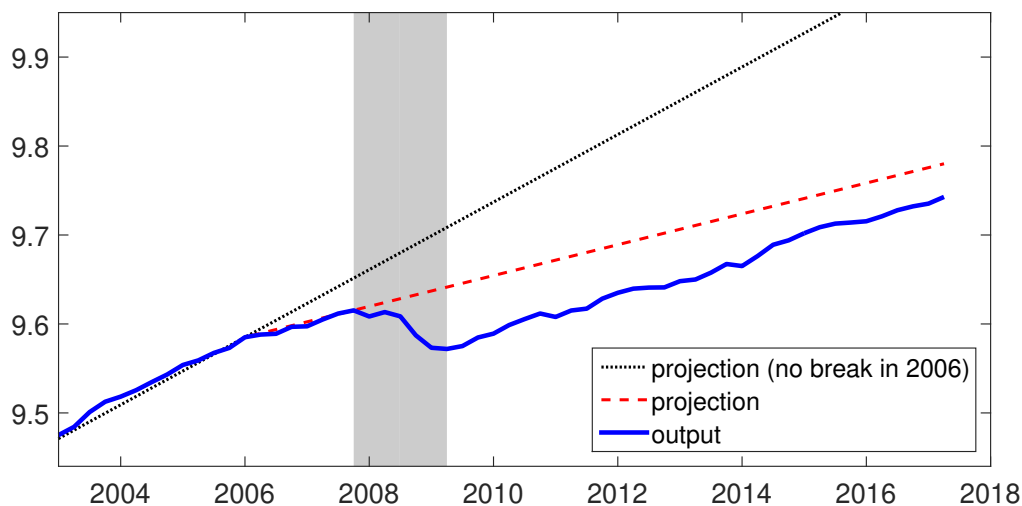
the same as the restricted version of a contractionary regime considered in [Hamilton \(1989\)](#). For the transition probabilities, q_1 and q_2 in (2) are estimated to be 0.029 and 0.033, respectively, and p and r are 0.735 and 0.802, respectively. The estimates of the transition probabilities suggest that expansions are more persistent (the continuation probability of the expansionary regime is $0.938=1-0.029-0.033$) than both types of recessions, much like the NBER reference cycle. The expected duration of the L-shaped regime is 3.78 quarters, while it is 5.05 quarters for the U-shaped regime.

3.1 Time-Varying Mean and Trend Projections

Figure 2 depicts the estimated time-varying mean from the benchmark specification using $E[\mu_t|I_t]$ where $\mu_t \equiv \Delta y_t - e_t$ and $I_t \equiv (\Delta y_1, \Delta y_2, \dots, \Delta y_t)$. Closely tracking realized real GDP growth and reflecting $\hat{\delta} = -0.52$, the time-varying mean declines in 2006, with this reduction in trend growth clearly contributing to the weak recovery of the U.S. economy following the Great Recession.⁴

⁴Figure 2 looks similar to and compares favorably with the estimated time-varying mean in [Eo and Kim \(2016\)](#) for a Markov-switching model with time-varying regime-dependent mean growth rates that depend on each other across booms and recessions and allows for possible structural change in trend growth. The point is that our simpler model also allows a structural break in trend growth and can capture differences in

Figure 3: Projected Trends in 2006:Q1 and Realized Output



Notes: We calculate trend projections in 2006:Q1 (the structural break date in output growth) (i) assuming no break with the dotted black line and (ii) accounting for the structural break in output growth with the dashed red line. The shaded area denotes the Great Recession.

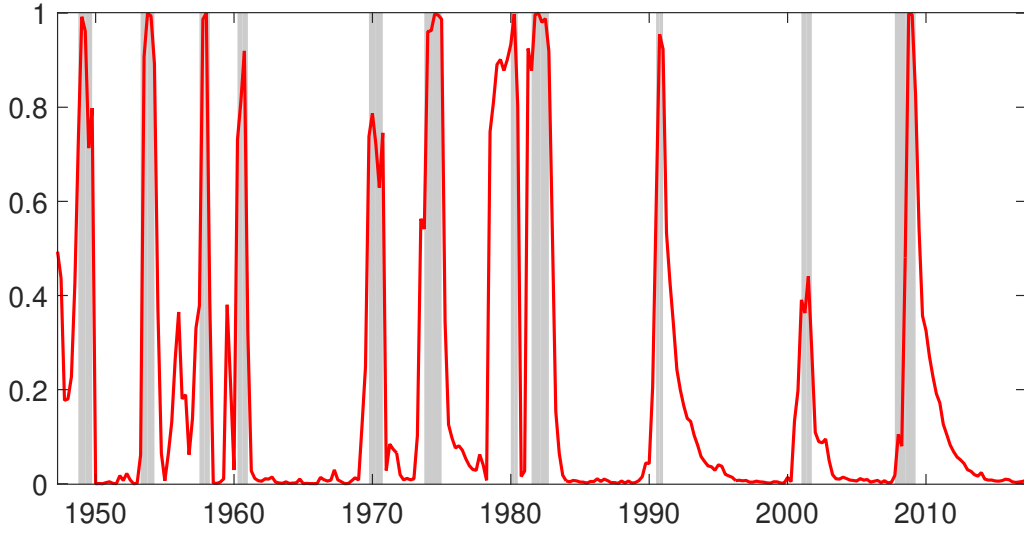
To demonstrate the magnitude of the trend break in 2006:Q1, we plot trend projections in Figure 3, both accounting for and not accounting for the structural break. The black dotted line shows the projection without the break using the pre-break expansionary mean growth rate of $\hat{\mu}_0 = 0.96$, which diverges markedly from realized output (solid blue line) even before the Great Recession. The red dashed line depicts the projection using the expansionary mean growth rate of $\hat{\mu}_0 + \hat{\delta} = 0.44$ after taking into account the decline in trend growth in 2006:Q1 and clearly supports the idea that the trend growth decline began in 2006 prior to the onset of the Great Recession.

3.2 Characterizing Recessions as L shaped or U shaped

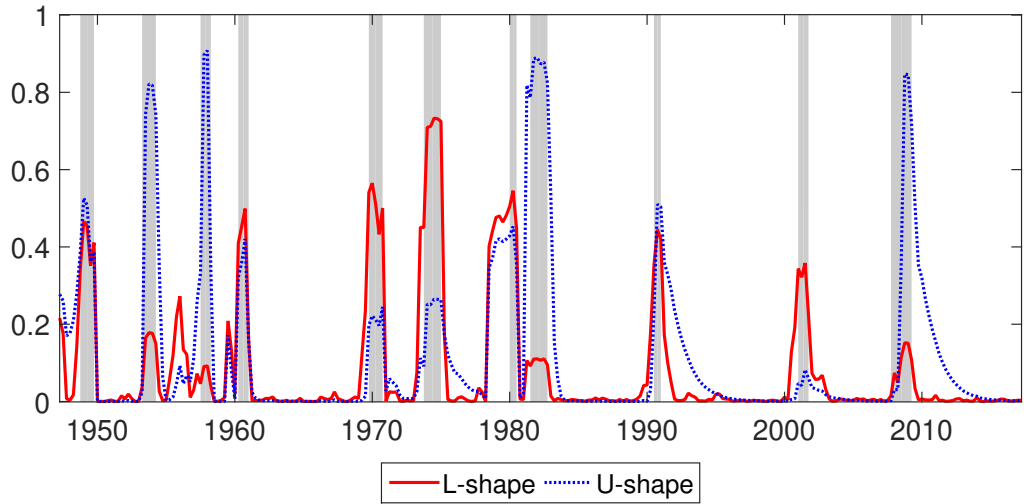
Figure 4(a) plots the smoothed probability of being in a contractionary regime in period t calculated from the sum of the probabilities of being in the L-shaped regime or the U-shaped regime using $E[t = contraction|I_T] \equiv E[S_t = 1|I_T] + E[S_t = 2|I_T]$. The probability closely

mean growth for each recession and expansion depending whether the contractionary regime that generates a recession is L or U shaped, with the mean growth in a recession related to the mean growth in the subsequent expansion and a long-run relationship between mean growth rates in recessions and expansions such that a long-run expectation of output growth exists. See [Eo and Kim \(2016\)](#) for more details.

Figure 4: Smoothed Probabilities of Contractionary Regimes



(a) Probability of a Contractionary Regime



(b) Probabilities of L-shaped and U-shaped Regimes

Notes: The probability of a contractionary regime is the sum of the probabilities of the L-shaped and U-shaped regimes. The shaded areas denote NBER recession dates.

matches the timing of NBER recessions. For nine of the eleven NBER recessions in the sample, the smoothed probability is greater than 90% throughout most of a given recession. The 1969-70 and 2001 recessions are the exceptions, with the probability close to 80% for the 1969-70 recession and slightly below 50% for the 2001 recession.

Figure 4(b) depicts the smoothed probability of being in an L-shaped regime with the

solid red line and that of being in the U-shaped regime with the dotted blue line. Considering the relative contribution of two different types of recession for each episode, the probabilities suggest that the 1969-70, 1973-75, and 2001 recessions can be characterized as L shaped, while the 1953-54, 1957-58, 1981-82, and 2007-09 recessions can be characterized as U shaped. The appropriate characterizations of the 1948-49, 1960-61, 1980, and 1990-91 recessions are less conclusive.

We find it notable that the probability of a U-shaped regime is significantly higher than that of an L-shaped regime during the 2007-09 recession. So the conventional view of the Great Recession as being L-shaped, perhaps due to its financial origins, is not supported by this model and the data.⁵ Instead, the lower level and growth of output since the Great Recession appears to be driven by a reduction in trend growth that began in 2006 prior to the onset of the recession in 2007:Q4.

At the same time, the probability of a U-shaped regime remained elevated after the trough date established by the NBER for the 2007-09 recession, as it also did for the 1990-91 recession. This could be related to a weak labor market (“jobless recovery”) after these recessions. Also, the zero-lower-bound on interest rates restricted the ability of monetary policy to help stimulate a strong recovery after the 2007-09 recession. Thus, the weak growth since the Great Recession could be partly related to a large and persistent negative output gap following the recession. We examine this possibility next.

3.3 Estimated Output Gap

An L-shaped regime has a permanent effect on output, while a U-shaped regime only has a transitory effect. This difference in the long-run effects of the two different types of contractionary regimes has important implications for trend output and the output gap.

We define the output gap as transitory deviations away from trend and use the Beveridge-Nelson (BN) decomposition ([Beveridge and Nelson \(1981\)](#)) as the basis for thinking about

⁵See, for example, [Cerra and Saxena \(2008\)](#), [Reinhart and Rogoff \(2009\)](#), and [Jordà, Schularick and Taylor \(2017\)](#), amongst many others, on the idea that financial crisis recessions, of which the Great Recession is a clear example, have large and persistent negative effects on the level of economic activity.

trend. In particular, the BN trend is

$$\hat{\tau}_t^{BN} = \lim_{h \rightarrow \infty} \left\{ E^M [y_{t+h} | I_t] - h \cdot E^M [\Delta y_t] \right\}, \quad (4)$$

where $E^M[\cdot]$ is the expectations operator with respect to a time series model and I_t is the set of relevant and available information observed up to time t . The BN trend is equivalent to the long-horizon conditional forecast of the level of output minus any deterministic drift. This implies that, as the forecasting horizon extends to infinity, a long-horizon forecast of output should no longer be influenced by the realization of the transitory component at time t .

Because we consider a Markov-switching time series model, we adopt the generalization of the BN decomposition for regime-switching processes developed in [Morley and Piger \(2008\)](#). This regime-dependent steady-state approach (RDSS) approach involves constructing long-horizon forecasts conditional on sequences of regimes and then marginalizing over the distribution of the unknown regimes. In particular, the RDSS trend is given by

$$\hat{\tau}_t^{RDSS} = \sum_{\tilde{S}_t} \left\{ \hat{\tau}_t^{RDSS}(\tilde{S}_t) \cdot p^M(\tilde{S}_t | I_t) \right\}, \quad (5)$$

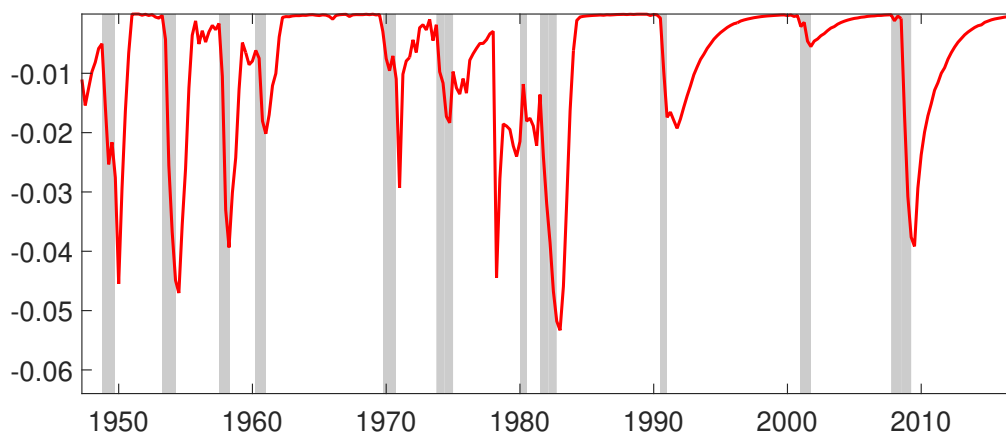
where

$$\hat{\tau}_t^{RDSS}(\tilde{S}_t) = \lim_{h \rightarrow \infty} \left\{ E^M \left[y_{t+h} | \{S_{t+k} = i^*\}_{k=1}^h, \tilde{S}_t, I_t \right] - h \cdot E^M \left[\Delta y_t | \{S_{t+k} = i^*\}_{-\infty}^\infty \right] \right\}, \quad (6)$$

$\tilde{S}_t = (S_t, \dots, S_{t-m})'$ is a vector of relevant current and past regimes for forecasting output, $p^M(\cdot)$ is a probability distribution with respect to the time series model M , S_t is the latent state variable in (1) that evolves according to a fixed transition matrix (2), and i^* is the “normal” regime in which the mean of the transitory component is assumed to be 0. Unlike the traditional BN decomposition, there is no implicit assumption that the cycle is unconditionally mean 0 and we choose the expansion regime as the normal regime $i^* = 0$.⁶ Furthermore, unlike the unobserved components models in [Kim and Murray \(2002\)](#) and

⁶See [Morley and Piger \(2008\)](#) for a full discussion of this choice and [Morley and Piger \(2012\)](#) for a justification of choosing the expansionary regime as the normal regime.

Figure 5: Estimated Output Gap for the Benchmark Specification



Note: The shaded areas denote NBER recession dates.

Kim and Piger (2002), there is no assumption that the underlying permanent and transitory shocks are uncorrelated (see Morley and Piger (2008)). Meanwhile, the probability $p^M(\cdot)$ can be evaluated via the Hamilton (1989) filter. Once $\hat{\tau}_t^{RDSS}$ in (5) is available, the estimated output gap, \hat{c}_t^{RDSS} , can be easily calculated as

$$\hat{c}_t^{RDSS} = y_t - \hat{\tau}_t^{RDSS}. \quad (7)$$

Figure 5 presents the estimated output gap for the benchmark specification in (3). We can see that large negative movements the output gap closely match up with NBER recessions. Because $\hat{\lambda}_1 \approx 0$ in Table 1, L-shaped recessions should affect trend output, not the output gap. Thus, the large negative movements in the output gap correspond primarily to the U-shaped recessions.

In terms of the Great Recession, the probability of contraction spikes up and the output gap declines later than the NBER peak date of 2007:Q4 in Figures 4 and 5. As Figure 3 makes clear, the reason for this delay is that the level of real GDP does not decline sharply until the second half of 2008, although real GDP did not grow at its usual expansionary rate in the first half of 2008, even accounting for the structural break in trend growth. The delayed timing of the severe contraction for the Great Recession is distinct from the behaviour of real GDP in previous recessions and possibly reflects a misattribution by the

NBER of a particularly lackluster manifestation of weak trend growth during the first half of 2008 to being part of the recession phase.⁷

4 Structural Change in Trend Growth

In our benchmark specification, we assumed a structural break in trend growth occurred in 2006:Q1. In this section, we confirm that the break in 2006:Q1 is supported by the data and plays the crucial role in understanding why the level and growth of U.S. real GDP were so much lower than projected prior to the Great Recession. We also demonstrate the importance of accounting for this structural break in order to maintain a good forecasting relationship between the output gap and inflation.

4.1 Existence and Timing of Structural Breaks in U.S. Real GDP

We first consider potential structural break dates for trend growth by applying [Qu and Perron's](#) (2007) testing procedure for multiple structural breaks in mean and variance to quarterly output growth for the sample period of 1947:Q2 to 2017:Q2 with 15% trimming at the beginning and the end of the sample period and between break dates.⁸ We find evidence of two breaks, which are estimated to have occurred in 1984:Q2 and 2006:Q1, as reported in [Table 2](#). These dates align with the timing of the so-called Great Moderation widely reported in the literature (e.g., [Kim and Nelson \(1999b\)](#) and [McConnell and Perez-Quiros \(2000\)](#)) and the break date for the reduction in output growth used in our benchmark specification that was also found in [Luo and Startz \(2014\)](#) and [Kamber, Morley and Wong \(2017\)](#). The structural breaks are significant at the 5% level and there is no support for an additional break even at a 10% level. Related to the Great Moderation and our Markov-switching model, we note that a larger variance for output growth before 1984:Q2 could potentially be related to a more frequent realization of recessions before the mid-1980s. In

⁷It may instead be related to an end-of-expansion overhiring phenomenon (see [Gordon \(2003\)](#)) that could have lowered productivity before the onset of an actual recession in the second half of 2008.

⁸The test regression includes only a constant. However, we allow for potential serial correlation in output growth, as would be implied by our Markov-switching model, by employing a heteroskedasticity and autocorrelation consistent (HAC) estimator of the long-run variance following [Andrews and Monahan \(1992\)](#).

Table 2: Structural Break Tests for Output Growth

# of breaks	Estimated Break Dates	LR Test Stat	Critical Value (5%)
1	1984:Q2	66.19	12.09
2	1984:Q2, 2006:Q1	22.82	13.39
3	1960:Q4, 1984:Q2, 2006:Q1	9.14	14.28

Table 3: Estimates for Mean and Standard Deviation of Output Growth Allowing for Structural Breaks

Regime	Estimated Break Date	Mean	Std. Dev.	Confidence Set for Break Date
(a) Unrestricted Model				
1		0.89	1.16	
2	1984:Q2	0.80	0.49	[1982:Q1,1987:Q1]
3	2006:Q1	0.35	0.62	[1994:Q4,2006:Q4]
(b) Restricted Model				
1		0.82	1.17	
2	1984:Q2	0.82	0.49	[1982:Q1,1987:Q2][1991:Q1]
3	2006:Q1	0.35	0.62	[1994:Q4,2006:Q4]

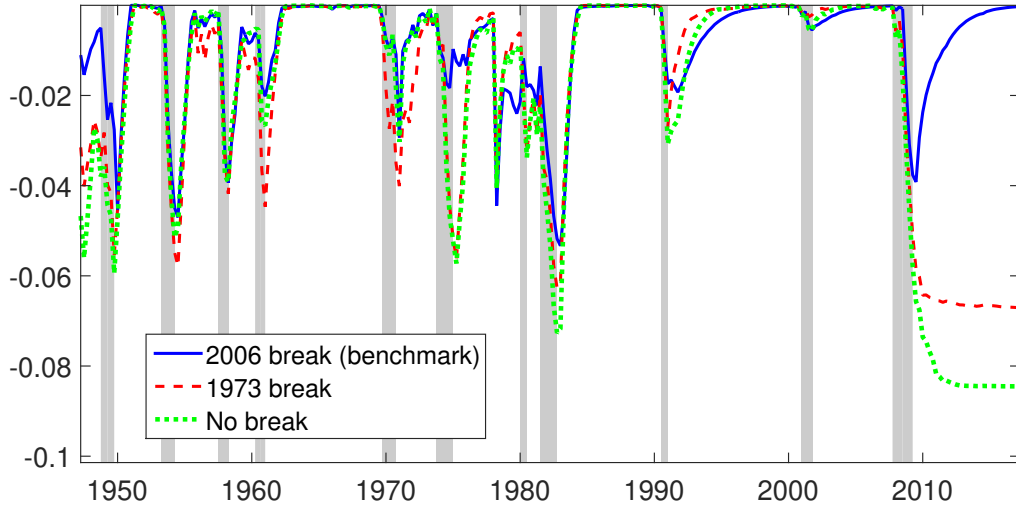
Note: The restricted model reported in panel (b) allows a change in variance only for the first break.

particular, the postwar U.S. economy experienced eight recessions between 1947 to 1984 (37 years), but only three recessions between 1985 to 2017 (33 years).

Estimates for the mean and standard deviation of output growth based on the estimated break dates, along with the confidence sets for the break dates, are reported in Table 3. The confidence set for the first break date covers a reasonably short interval of 1982:Q1 to 1987:Q1, while the confidence set for the second break date is much wider and ranges from 1994:Q4 to 2006:Q4.⁹ The estimated break date of 2006:Q1 is consistent with the date for the growth shortfall in Fernald et al. (2017) and they argue that it reflects slow growth of total factor productivity and a decline in labor force participation that are unrelated to the financial crisis and the Great Recession.

⁹We calculate the confidence sets using the inverted likelihood-ratio test approach in Eo and Morley (2015). For more details and comparison to other approaches for calculating confidence intervals/sets for structural break dates, see Eo and Morley (2015). We note that 2006:Q4 represents the last possible break date given 15% trimming and the confidence set would extend to the last possible break date of 2010Q2 given 10% trimming. However, 2006:Q1 remains the estimated break date given 10% trimming.

Figure 6: Estimated Output Gaps for Different Structural Break Assumptions



Note: The shaded areas denote NBER recession dates.

For the first estimated break in 1984:Q2, a likelihood ratio test of no change in mean suggest that the break corresponds to a change in variance only, with the sample standard deviation of output growth dropping by more than 50%. Note that the average growth rates before and after the first break date of 1984:Q2 are 0.89 and 0.80, respectively, and are very close as compared to the average growth rate of 0.35 after the second break date of 2006:Q1. Thus, we also report the results from the restricted model used in the likelihood ratio test that assumes mean growth does not change following the first break date of 1984:Q2, but only after the second break date of 2006:Q1. The parameter estimates and confidence sets for the break dates are very similar to those for the unrestricted model. The average growth rate declines by 0.47 in 2006:Q1, which is close to $\hat{\delta} = -0.52$ for the reduction in trend growth for the benchmark specification of our Markov-switching model reported in Table 1.

4.2 Output Gap and Trend Output under Different Assumptions about Structural Breaks

Figure 6 plots the estimated output gaps implied by our model under different assumptions about a structural break in trend growth. In addition to our benchmark specification of a

Table 4: Maximum Likelihood Estimates under Alternative Structural Break Assumptions

Parameter	1973 Break		No Break	
	Estimate	S.E.	Estimate	S.E.
q_1	0.0038	(0.0043)	0.0069	(0.0067)
q_2	0.0445	(0.0171)	0.0420	(0.0172)
p	0.9906	(0.0150)	0.9896	(0.0141)
r	0.6985	(0.1063)	0.6927	(0.1203)
σ^2	0.4744	(0.0468)	0.4931	(0.0487)
μ_0	0.9826	(0.0623)	0.8259	(0.0470)
μ_1	-2.0951	(0.4781)	-2.6951	(0.4634)
λ_1	0.3204	(0.0839)	0.4025	(0.0773)
μ_2	-1.8676	(0.1759)	-1.7743	(0.2291)
δ	-0.2599	(0.0854)		
log-lik	-343.88		-347.78	

Note: The 1973 break specification allows for a reduction in trend growth in 1973:Q1 in (3). The standard errors of the parameter estimates are reported in parentheses. The sample period is 1947:Q2 to 2017:Q2.

break in 2006:Q1, we also consider a break in 1973:Q1, consistent with the “productivity growth slowdown” in the early 1970s (e.g. Perron (1989)), and the case of no break. Since the Great Recession, the average growth rate of U.S real GDP has been 0.53 for the period of 2009:Q3 to 2017:Q2. In the absence of allowing for a break in mean in 2006 for the 1973 break specification or the no break specification, the estimated output gap remains persistently large and negative until the end of the sample in order to capture the weak output growth over this period, although the estimated output gaps are reasonably similar in all three cases prior to 2006.

Table 4 presents maximum likelihood estimates for the 1973 break specification and the no break specification of our Markov-switching model. The mean growth rate in the expansionary regime before the Great Recession is estimated to be $\hat{\mu}_0 + \hat{\delta} = 0.72$ for the 1973 break specification and $\hat{\mu}_0 = 0.83$ for no break specification. The differences between the model-implied expansionary growth rates and the average quarterly output growth rate since the Great Recession are then 0.19 for the 1973 break specification and 0.30 for no break specification, respectively. For these specifications, the L-shaped regime has a significant bounce-back effect ($\hat{\lambda}_1 = 0.32$ and 0.40, respectively) and the continuation of this regime

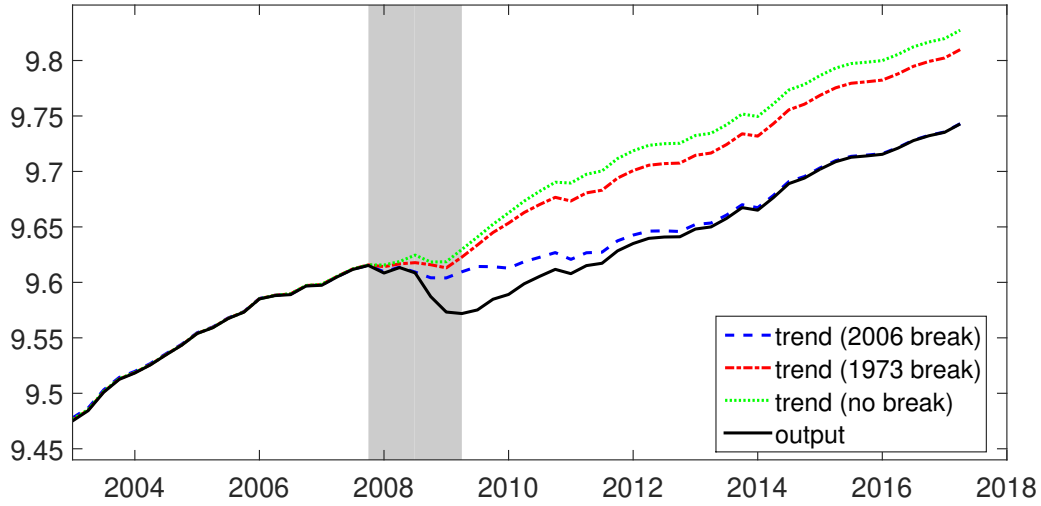
from the Great Recession until the end of the sample helps account for the large discrepancy in growth rates. Once taking into account the bounce-back effect for the L-shaped recession, the estimated mean growth rates after 6 quarters during the L-shaped regime are $\hat{\mu}_0 + \hat{\delta} + \hat{\mu}_1 + 6\hat{\lambda}_1 = 0.55$ (after the break) for the 1973 break specification and $\hat{\mu}_0 + \hat{\mu}_1 + 6\hat{\lambda}_1 = 0.55$ for the no break specification.¹⁰ These values are consistent with the average growth rate of output since the Great Recession. Because we find the L-shaped regime persisted from 2008:Q2 to the end of the sample for these specifications, the estimated continuation probability for the L-shaped regime for both specifications is extremely high with the value of 0.99, while that for the U-shaped regime is only 0.69.¹¹ Notably, the likelihood values for these alternative specifications are lower than for the benchmark specification, although the distributions of likelihood ratio tests for a structural break or a given break date are not known in the context of a Markov-switching process. Thus, we rely on the results in Tables 2 and 3 to justify the existence of a structural break in trend growth and to use 2006:Q1 as the break date instead of 1973:Q1 for our benchmark specification.

Figure 7 plots trend output from (5) around the Great Recession for the different specifications of a possible structural break in trend growth. Realized output has remained well below the estimated trends since the Great Recession began for the 1973 break specification and the no break specification, but the estimated trend for the benchmark specification with a break in 2006 implies that the U.S. economy fully recovered from the Great Recession by 2014. Thus, the relatively lower level and growth of output since the Great Recession appear to be driven by a reduction in trend growth that began prior to the onset of the recession rather than the large and persistent negative output gap after the recession or permanent effects of the recession itself.

¹⁰The L-shaped regime still has a permanent, negative effect on the level of output because $\hat{\mu}_1 + 6\hat{\lambda}_1 < 0$ for both specifications. Hence, with reference back to Figure 1, it is still reasonable to label these as L-shaped regimes.

¹¹Thus, the L-shaped regime state variable effectively becomes a structural break dummy for these alternative specifications. Note that the continuation probabilities for the L-shaped and U-shaped recessions for the benchmark specification have more similar values of 0.74 and 0.80, respectively.

Figure 7: Output and Trend for different break dates around the Great Recession



4.3 Forecasting Inflation

We also examine the importance of the timing of the structural break by conducting an out-of-sample forecasting comparison for inflation with a Phillips Curve type forecasting equation and the different estimated output gaps. Following [Clark and McCracken \(2006\)](#) we specify an autoregressive distributed lag (ADL) forecasting model which is fairly standard in the forecasting literature (e.g. [Stock and Watson \(1999, 2009\)](#)). For an h -period-ahead inflation forecast, the ADL model is given by

$$\pi_{t+h} - \pi_t = \alpha + \sum_{j=0}^{p-1} \phi_j \Delta \pi_{t-j} + \kappa \hat{c}_t^{RDSS} + \epsilon_{t+h,t}, \quad (8)$$

where π_t is inflation and \hat{c}_t^{RDSS} is the estimated output gap in (7) that depends on the structural break specification. Inflation is calculated as 400 times the first difference of the natural logarithms of the headline and core measure of Personal Consumption Expenditures (PCE) price index for the sample period of 1959:Q2 to 2017:Q2.¹² We consider the core PCE inflation rate in addition to the headline PCE inflation rate because the Federal Reserve tends

¹² The raw data for inflation are seasonally adjusted quarterly Personal Consumption Expenditures: Chain-type Price Index and Personal Consumption Expenditures: Chain-type Price Index Excluding Food and Energy were obtained from the St. Louis Fed (FRED) database. We use the data from 1959:Q2 because of availability for core PCE index.

Table 5: Inflation Forecast Evaluation

	Headline PCE Inflation							
	h=1		h=2		h=3		h=4	
	RRMSE	DM	RRMSE	DM	RRMSE	DM	RRMSE	DM
1973 Break	1.15	0.08	1.30	0.04	1.24	0.09	1.32	0.11
No Break	1.23	0.03	1.47	0.02	1.41	0.04	1.61	0.06
	Core PCE Inflation							
	h=1		h=2		h=3		h=4	
	RRMSE	DM	RRMSE	DM	RRMSE	DM	RRMSE	DM
1973 Break	1.71	0.00	2.14	0.00	2.53	0.01	2.55	0.04
No Break	1.97	0.00	2.60	0.00	3.09	0.01	3.13	0.03

Note: RRMSE reports the ratio of the root mean squared error of the forecasts for a model using a particular estimated output gap compared to that for a model using the estimated output gap from our benchmark specification with a break in trend growth in 2006:Q1. DM reports Diebold-Mario test p -values calculated under the null of the equality of forecast accuracy for a two-sided test.

to focus on it as the key measure for tracking underlying inflationary trends.¹³ We calculate out-of-sample forecasts over the evaluation sample of 2009:Q3 to 2017:Q2 using parameter estimates based on data up until 2009:Q2, with the lag order p for inflation chosen based on AIC. We find that the ADL model for headline inflation includes two lags ($p = 2$), while the ADL model for core inflation includes one lag ($p = 1$).

The results in Table 5 show that the estimated output gap for our benchmark specification outperforms the other estimates at all horizons ($h = 1, \dots, 4$ quarters) in terms of relative root mean squared error (RRMSE) and always significantly so based on Diebold-Mariano (DM) tests (Diebold and Mariano, 2002) in the case of core inflation. Strikingly, the RRMSEs are often greater than 2 in the case of core inflation.

These results suggest it may not be necessary to allow for a change in the slope of the Phillips Curve, as discussed in Ball and Mazumder (2011), to capture the behavior of inflation since the Great Recession. However, it is clearly important to have an estimated output gap that does not remain persistently negative up until the end of the sample to accurately forecast inflation. By allowing for a structural break in trend growth in 2006:Q1, the benchmark specification of our Markov-switching model produces such an output gap.

¹³See, for example, Federal Reserve Board of Governors (2004) and Bernanke (2007).

In particular, given a full recovery from the Great Recession by 2014, the estimated output gap for our benchmark specification does not imply that the economy should be experiencing deflation. These results are consistent with [Huang and Luo \(2017\)](#), who find that estimates of the output gap that allow for a lower level of trend since the Great Recession can explain the behavior of inflation in recent years.

5 Conclusion

Perhaps surprisingly, we find that the Great Recession was U shaped. Thus, the recession does not, in itself, explain the stagnation of U.S. real GDP since it ended. It did generate a large and persistent negative output gap. However, the U.S. economy full recovered from it by 2014. Instead, the low output and growth since the Great Recession appears to be due to a secular decline in trend growth that began in 2006 prior to the onset of the Great Recession. This finding is consistent with [Fernald et al. \(2017\)](#) and argues against the idea of hysteresis and super-hysteresis explaining economic stagnation since the Great Recession.

Notably, our results are based on a new highly flexible nonlinear time series model that we develop for the purpose of comparing the different explanations for why the level and growth of U.S. real GDP have been much lower than would have been projected prior to the Great Recession. Our Markov-switching model allows a given recession and its recovery to be either L or U shaped and fits the U.S. data well, with a clear characterization of most postwar recessions as either L or U shaped. The model also implies a highly asymmetric output gap that has a relatively good forecasting relationship with inflation. Thus, it may not be necessary to assume a change in the slope of the Phillips Curve to understand the behaviour of inflation since the Great Recession. However, it is necessary to account for the structural break in trend growth in 2006.

References

- Andrews, Donald W K, and J Christopher Monahan. 1992. “An Improved Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimator.” *Econometrica*, 60(4): 953–66.
- Antolin-Diaz, Juan, Thomas Drechsel, and Ivan Petrella. 2017. “Tracking the slowdown in long-run GDP growth.” *Review of Economics and Statistics*.
- Ball, Laurence. 2014. “Long-term damage from the Great Recession in OECD countries.” *European Journal of Economics and Economic Policies: Intervention*, 11(2): 149–160.
- Ball, Laurence, and Sandeep Mazumder. 2011. “Inflation Dynamics and the Great Recession.” *Brookings Papers on Economic Activity*, 337.
- Bernanke, Ben S. 2007. “Federal Reserve Communications.”
- Beveridge, Stephen, and Charles R Nelson. 1981. “A new approach to decomposition of economic time series into permanent and transitory components with particular attention to measurement of the business cycle.” *Journal of Monetary economics*, 7(2): 151–174.
- Blanchard, Olivier, Eugenio Cerutti, and Lawrence Summers. 2015. “Inflation and activity—Two explorations and their monetary policy implications.” National Bureau of Economic Research.
- Blanchard, Olivier J, and Lawrence H Summers. 1986. “Hysteresis and the European unemployment problem.” *NBER macroeconomics annual*, 1: 15–78.
- Carrasco, Marine, Liang Hu, and Werner Ploberger. 2014. “Optimal test for Markov switching parameters.” *Econometrica*, 82(2): 765–784.
- Cerra, Valerie, and Sweta Chaman Saxena. 2008. “Growth dynamics: the myth of economic recovery.” *American Economic Review*, 98(1): 439–57.
- Cerra, Valerie, and Sweta Chaman Saxena. 2017. “Booms, Crises, and Recoveries: A New Paradigm of the Business Cycle and its Policy Implications.” International Monetary Fund.
- Clark, Todd E, and Michael W McCracken. 2006. “The predictive content of the output gap for inflation: Resolving in-sample and out-of-sample evidence.” *Journal of money, credit, and Banking*, 38(5): 1127–1148.
- Diebold, Francis X, and Robert S Mariano. 2002. “Comparing predictive accuracy.” *Journal of Business & economic statistics*, 20(1): 134–144.

- Eo, Yunjong, and Chang-Jin Kim. 2016. “Markov-Switching Models with Evolving Regime-Specific Parameters: Are Postwar Booms or Recessions All Alike?” *Review of Economics and Statistics*, 98(5): 940–949.
- Eo, Yunjong, and James Morley. 2015. “Likelihood-ratio-based confidence sets for the timing of structural breaks.” *Quantitative Economics*, 6(2): 463–497.
- Federal Reserve Board of Governors. 2004. “Monetary Report to the Congress.”
- Fernald, John G. 2015. “Productivity and Potential Output before, during, and after the Great Recession.” *NBER Macroeconomics Annual*, 29(1): 1–51.
- Fernald, John G, Robert E Hall, James H Stock, and Mark W Watson. 2017. “The Disappointing Recovery of Output after 2009.” *Brookings Papers on Economic Activity*.
- Fernald, John G, et al. 2016. “Reassessing longer-run US growth: how low?” *Federal Reserve Bank of San Francisco Working Paper*, 18.
- Gordon, Robert J. 2003. “Exploding productivity growth: context, causes, and implications.” *Brookings Papers on Economic Activity*, 2003(2): 207–298.
- Gordon, Robert J. 2015. “Secular Stagnation: A Supply-Side View.” *American Economic Review*, 105(5): 54–59.
- Grant, Angelia L, and Joshua CC Chan. 2017. “A Bayesian Model Comparison for Trend-Cycle Decompositions of Output.” *Journal of Money, Credit and Banking*, 49(2-3): 525–552.
- Hamilton, James D. 1989. “A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle.” *Econometrica*, 57(2): 357–84.
- Hansen, Alvin H. 1939. “Economic progress and declining population growth.” *The American Economic Review*, 29(1): 1–15.
- Huang, Yu-Fan, and Sui Luo. 2017. “Potential output and inflation dynamics after the Great Recession.” *Empirical Economics*, 1–23.
- Huang, Yu-Fan, Sui Luo, and Richard Startz. 2016. “Are recoveries all the same: GDP and TFP?” *Manuscript, University of California Santa Barbara*.
- Jordà, Òscar, Moritz Schularick, and Alan M Taylor. 2017. “Macrofinancial history and the new business cycle facts.” *NBER Macroeconomics Annual*, 31(1): 213–263.

- Kamber, Gunes, James Morley, and Benjamin Wong. 2017. “Intuitive and Reliable Estimates of the Output Gap from a Beveridge-Nelson Filter.” *Review of Economics and Statistics*, forthcoming.
- Kim, Chang-Jin, and Charles R. Nelson. 1999a. “Friedman’s Plucking Model of Business Fluctuations: Tests and Estimates of Permanent and Transitory Components.” *Journal of Money, Credit and Banking*, 31(3): 317–34.
- Kim, Chang-Jin, and Charles R. Nelson. 1999b. “Has The U.S. Economy Become More Stable? A Bayesian Approach Based On A Markov-Switching Model Of The Business Cycle.” *Review of Economics and Statistics*, 81(4): 608–616.
- Kim, Chang-Jin, and Christian J Murray. 2002. “Permanent and transitory components of recessions.” In *Advances in Markov-Switching Models*. 19–39. Springer.
- Kim, Chang-Jin, and Jeremy Piger. 2002. “Common stochastic trends, common cycles, and asymmetry in economic fluctuations.” *Journal of Monetary Economics*, 49(6): 1189–1211.
- Kim, Chang-Jin, James Morley, and Jeremy Piger. 2005. “Nonlinearity and the permanent effects of recessions.” *Journal of Applied Econometrics*, 20(2): 291–309.
- Kim, Chang-Jin, Jeremy M Piger, and Richard Startz. 2007. “The dynamic relationship between permanent and transitory components of US business cycles.” *Journal of Money, Credit and Banking*, 39(1): 187–204.
- Luo, Sui, and Richard Startz. 2014. “Is it one break or ongoing permanent shocks that explains US real GDP?” *Journal of Monetary Economics*, 66: 155–163.
- McConnell, Margaret M., and Gabriel Perez-Quiros. 2000. “Output Fluctuations in the United States: What Has Changed since the Early 1980’s?” *American Economic Review*, 90(5): 1464–1476.
- Morley, James, and Jeremy Piger. 2008. “Trend/cycle decomposition of regime-switching processes.” *Journal of Econometrics*, 146(2): 220–226.
- Morley, James, and Jeremy Piger. 2012. “The asymmetric business cycle.” *Review of Economics and Statistics*, 94(1): 208–221.
- Morley, James C, Charles R Nelson, and Eric Zivot. 2003. “Why are the Beveridge-Nelson and unobserved-components decompositions of GDP so different?” *The Review of Economics and Statistics*, 85(2): 235–243.
- Perron, Pierre. 1989. “The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis.” *Econometrica*, 57(6): 1361–1401.

- Qu, Zhongjun, and Pierre Perron. 2007. “Estimating and Testing Structural Changes in Multivariate Regressions.” *Econometrica*, 75(2): 459–502.
- Reinhart, Carmen M, and Kenneth S Rogoff. 2009. “The aftermath of financial crises.” *American Economic Review*, 99(2): 466–72.
- Stock, James H, and Mark W Watson. 1999. “Forecasting inflation.” *Journal of Monetary Economics*, 44(2): 293–335.
- Stock, James H, and Mark W Watson. 2009. “Phillips curve inflation forecasts.” In *Understanding inflation and the implications for monetary policy*, ed. Jane Little Jeffrey Fuhrer, Yolanda Kodrzycki and Giovanni Olivei. MIT Press, Cambridge.
- Summers, Lawrence H. 2014. “US economic prospects: Secular stagnation, hysteresis, and the zero lower bound.” *Business Economics*, 49(2): 65–73.
- Summers, Lawrence H. 2015. “Demand side secular stagnation.” *The American Economic Review*, 105(5): 60–65.