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

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Abstract

Related to the idea of secular stagnation, the path of U.S. real GDP has been much lower than would have been projected prior to the Great Recession. We investigate the extent to which this apparent stagnation is due to hysteresis effects of the Great Recession, a large and persistent negative output gap following the recession, or slower trend growth. To do so, we develop a new Markov-switching time series model that allows a given recession to either have a large permanent effect or lead to a full recovery and also accounts for possible structural breaks in trend growth. Estimates for our model suggest that the Great Recession generated a negative and persistent output gap, but no permanent effects, with the economy fully recovering to trend by 2014. Instead, the relatively low level and growth path of output that continues after the recovery appears to be driven by a reduction in trend growth that began in 2006, prior to the onset of the Great Recession. Our results about the timing of the reduction in trend growth are supported by data on aggregate consumption and final sales, while our findings in terms of the role of the Great Recession in explaining secular stagnation help to reconcile the lack of deflation in recent years.

Keywords: Secular stagnation; Great Recession; output gap; Markov switching; structural breaks

JEL classification: C22; C51; E32; E37

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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. According to 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 path 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 paths, 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 started before the Great Recession (Fernald, 2015, 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 path of U.S. real GDP has been much lower in terms of both level and growth than would have been projected prior to the Great Recession. In particular, we investigate whether this apparent stagnation is due to hysteresis effects, a large and persistent negative output gap following the recession, or slower trend growth. Building on Hamilton (1989), Kim and Nelson (1999*a*), Kim, Morley and Piger (2005), and Eo and Kim (2016), our univariate Markov-switching model of output growth allows a given recession to either have a large permanent effect on the level of output (an "L-shaped" recession) or lead to a full recovery (a "U-shaped" recession). We also investigate possible structural breaks in trend growth. In particular, using the testing procedure in Qu and Perron (2007) and also considering data on aggregate consumption and final sales, our empirical analysis supports a reduction in trend growth in 2006. Allowing for this break in our Markov-switching model, we find that the Great Recession was U shaped, generating a negative and persistent output gap, but no hysteresis effects, with the economy fully recovering to trend 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 can also help explain a lack of deflation in recent years.

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), Kim and Nelson (1999*a*), 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 recessions is not predetermined 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 for L- and U-shaped recessions. However, our model is univariate and much more parsimonious in terms of capturing the main propagation of recessionary shocks. 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 both 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. 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 are also able to catagorize different 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, including for data on aggregate consumption and final sales, 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 also 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) a "hysteresis" regime with permanent L-shaped effects on the level of output, as in Hamilton (1989), and (ii) a "plucking" regime with only transitory U-shaped effects, as in Kim and Nelson (1999*a*). This structure of allowing a given recession to have either L- or U-shaped effects 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:

$$\Delta y_t = \mu_0 + \mu_1 \mathbf{1}(S_t = 1) + \mu_2 \mathbf{1}(S_t = 2) + \lambda \sum_{k=1}^{m} \mathbf{1}(S_{t-k} = 2) + e_t, \tag{1}$$

where $e_t \sim i.i.d.N(0, \sigma^2)$, 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 "hysteresis" contractionary regime, $S_t = 2$ for the "plucking" contractionary regime according to transition probabilities $Pr[S_t = j | S_{t-1} = i] = p_{ij}$ 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]

To identify the contractionary regimes as being associated with two types of recessions, we assume that the economy does not switch between different contractionary regimes without going through an expansionary regime first. This sequencing of regimes can be imposed using restrictions on the regime transition probabilities as follows: $p_{12} = 0$ for the "hysteresis" regime to "plucking" regime transition and $p_{21} = 0$ for the "plucking" regime to "hysteresis"

[1] Specifically, 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.

regime transition. Thus, the overall regime transition matrix is given by

$$\Pi = \begin{bmatrix} 1 - p_{01} - p_{02} & 1 - p_{11} & 1 - p_{22} \\ p_{01} & p_{11} & 0 \\ p_{02} & 0 & p_{22} \end{bmatrix}. \tag{2}$$

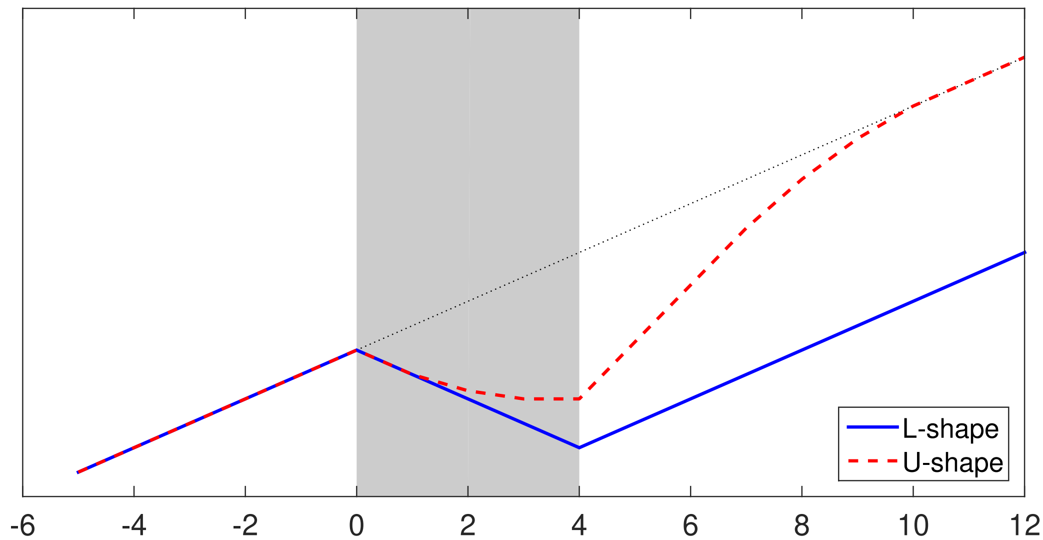
The λ parameter is the key distinct feature of the "plucking" contractionary regime in (1) because it allows for a bounce-back effect, as in Kim, Morley and Piger (2005). In particular, to clearly identify this contractionary regime as only having transitory U-shaped effects, as in Kim and Nelson (1999*a*), we impose the restriction $\mu_2 + m \cdot \lambda = 0$. Thus, the bounce-back effect $m \cdot \lambda$ exactly cancels out the contractionary effect μ_2.[2] Thus, the "hysteresis" contractionary regime is the only one that can have L-shaped permanent effects on the level of output.[3]

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 effect to $m = 6$ periods and the model parameters to be $\mu_0 = 1$ for the expansionary regime, $\mu_1 = -2$ for the L-shaped regime, and $\mu_2 = -2$ (implying $\lambda = 1/3$) 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 for 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 "plucking" regime, the bounce-back term takes effect as the recession continues and flattens out the path of output, with the economy then 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,

[2] In 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 allow for possible permanent effects with their U-shaped regime in addition to assuming permanent effects with the L-shaped regime.

[3] Typically 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 turns out to be identified given the restriction on λ and the transition probabilities only. Thus, we place no restrictions on the other conditional mean parameters in (1).

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



Note: The shaded area denotes the contractionary regime.

for the "hysteresis" 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:Q4 and were obtained from the St. Louis Fed database (FRED). 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+1} states in each period. The length of the post-recession bounce-back effect is set to $m = 6$ quarters following Kim, Morley and Piger (2005). Standard errors are based on numerical second derivatives.

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

Table 1: Maximum Likelihood Estimates for the Benchmark Specification

Parameter	Estimate	S.E.
p_{01}	0.0268	(0.0193)
p_{02}	0.0316	(0.0150)
p_{11}	0.7397	(0.1244)
p_{22}	0.8012	(0.0870)
σ^2	0.4367	(0.0463)
μ_0	0.9501	(0.0626)
μ_1	-1.1753	(0.2423)
μ_2	-1.9693	(0.1775)
λ	0.3282	(0.0296)
δ	-0.5000	(0.1264)
log-lik	-343.78	

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:Q4. We report estimates for both μ_2 and λ, although one implies the other given the restriction $\mu_2 + m \cdot \lambda = 0$.

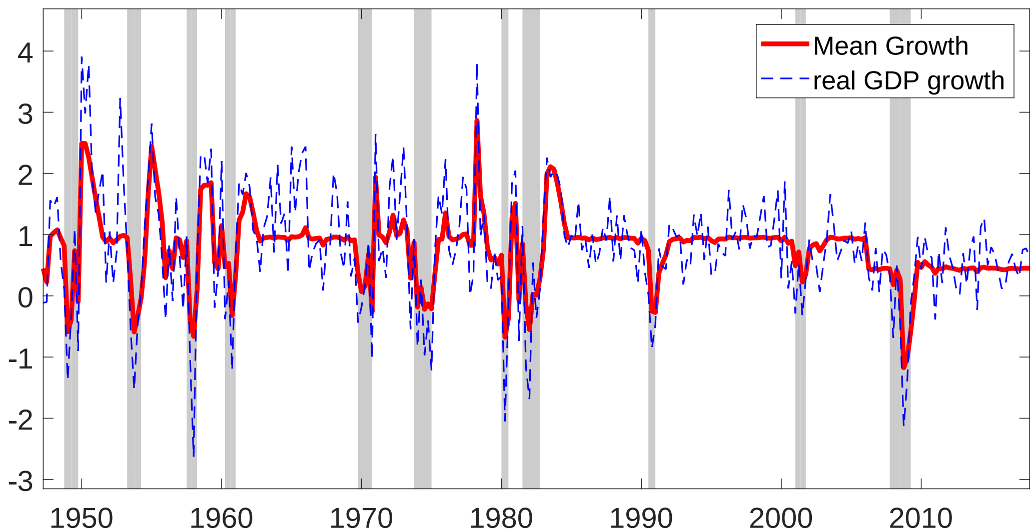
possibilities for the timing of a slowdown in more detail in Section 4. To account for the structural break, the model specification in (1) is altered accordingly:

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

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

Table 1 reports maximum likelihood estimates for the benchmark specification. The estimated output growth rates $\hat{\mu}_0 + \hat{\mu}_1 < 0$ for the "hysteresis" regime and $\hat{\mu}_0 + \hat{\mu}_2 < 0$ for the "plucking" regime indicate that both regimes are clearly contractionary, although this was not imposed in estimation. The estimates of the transition probabilities suggest that expansions are more persistent than both types of recessions, much like the NBER reference cycle. In particular, the implied continuation probability of the expansionary regime is 0.941=1-0.027-0.032, with expected duration of 16.95 quarters, while the expected duration is 3.84 quarters for the "hysteresis" regime and 5.03 quarters for the "plucking" regime.

Figure 2: Quarterly Output Growth and Time-Varying Mean



Note: The shaded areas denote NBER recession dates.

Meanwhile, the structural break in 2006:Q1 reduced trend growth by half a percentage point at a quarterly rate.

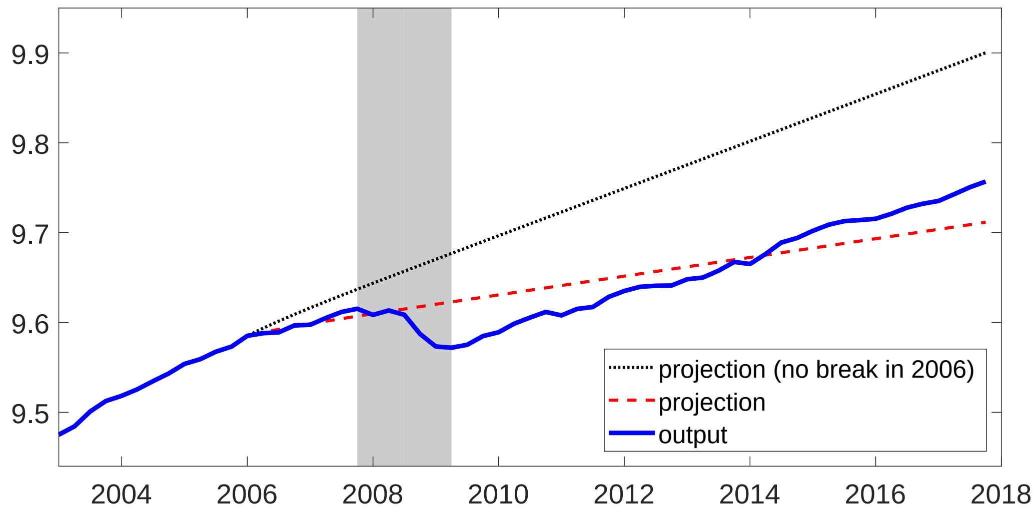
3.1 Time-Varying Mean and Output Projections

Figure 2 plots the estimated time-varying mean from the benchmark specification using $E[\mu_t|\Omega_t]$ where $\mu_t \equiv \Delta y_t - e_t$ and $\Omega_t \equiv (\Delta y_1, \Delta y_2, ..., \Delta y_t)$. Closely tracking realized real GDP growth and reflecting $\hat{\delta} = -0.50$, 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]

To demonstrate the magnitude of the trend break in 2006:Q1, Figure 3 plots projections from t = 2006:Q1 of future log output $E[y_{t+h}|\Omega_t]$, $h > 0$, both accounting for and not accounting for the structural break. The black dotted line shows the projection of log output

[4]Figure 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 mean growth for each recession and expansion based on 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 3: Output Projections in 2006:Q1 and Realized Output



Notes: Output and projections are reported in natural logs. We calculate 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.

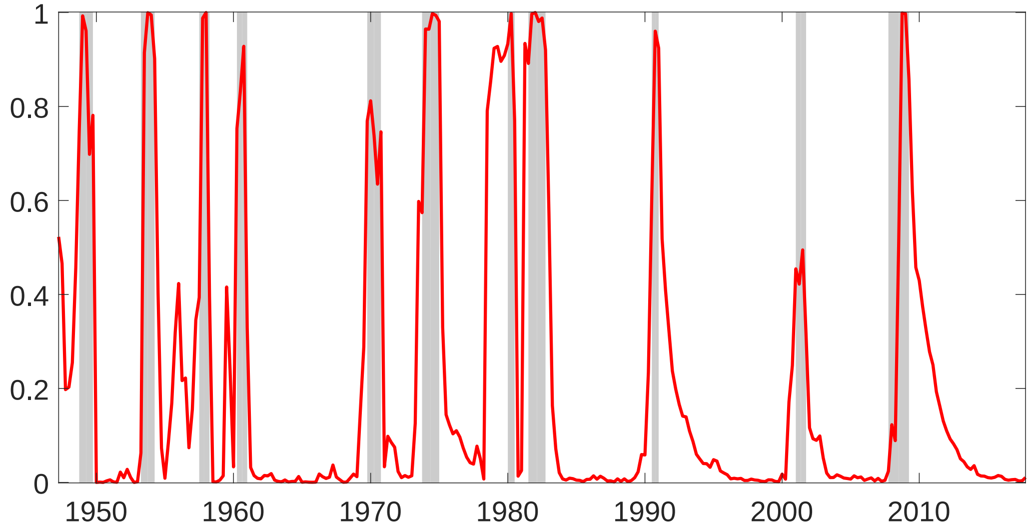
based without accounting for the structural break, which diverges markedly from realized output (solid blue line) even before the Great Recession. The red dashed line shows the projection accounting for the structural break and clearly supports the idea that the trend growth decline began in 2006 prior to the onset of the Great Recession.

3.2 Characterizing Recessions as L shaped or U shaped

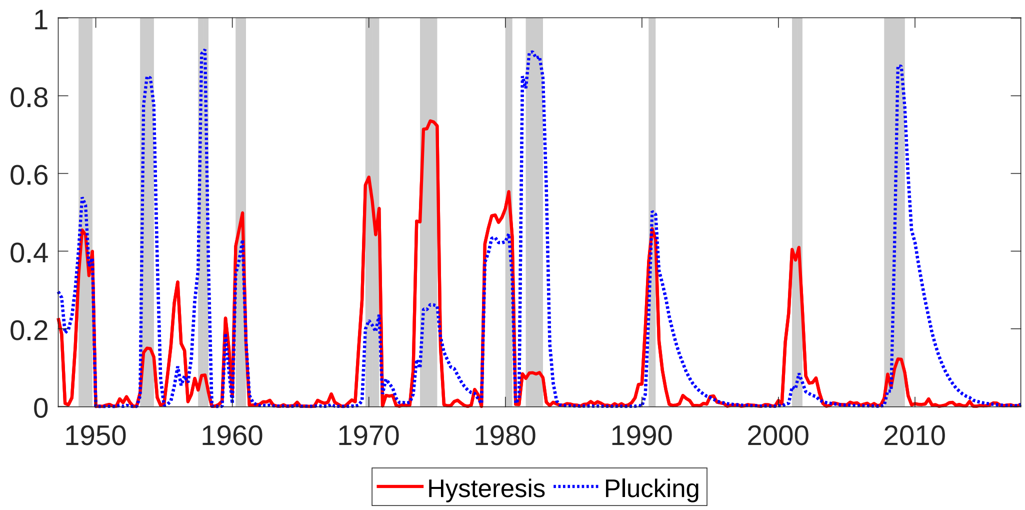
The top panel of Figure 4 plots the smoothed probability of being in a contractionary regime in period t calculated from the sum of the probabilities of being in the "hysteresis" regime or the "plucking" regime using $Pr[t = contraction|\Omega_T] \equiv Pr[S_t = 1|\Omega_T] + Pr[S_t = 2|\Omega_T]$. The probability closely 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.

The bottom panel of Figure 4 plots the underlying smoothed probabilities of the "hysteresis" regime and the "plucking" regime. Considering their relative contribution to the overall

Figure 4: Smoothed Probabilities of Contractionary Regimes



(a) Probability of a Contractionary Regime



(b) Probabilities of “Hysteresis” and “Plucking” Regimes

Notes: The probability of a contractionary regime is the sum of the probabilities of the “hysteresis” and “plucking” regimes. The shaded areas denote NBER recession dates.

probability of a contractionary regime, these 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 corresponding to the “plucking” regime is significantly higher than that of "hysteresis" regime during the 2007-09 period. In particular, the conventional view of the Great Recession as being L shaped, perhaps due to its financial origins, is not supported by the data with this model.[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 according to the NBER.

At the same time, the probability corresponding to the "plucking" 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 prolonged weak labor market ("jobless recovery") following these two recessions. Also, the zero-lower-bound on interest rates restricted the ability of monetary policy to help stimulate a strong recovery in the latter case. Thus, the weak growth following the Great Recession could be partly related to a large and persistent negative output gap after the recession. We examine this possibility next.

3.3 Estimated Output Gap

A "hysteresis" regime has a permanent effect on the level of output, while a "plucking" 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 trend. In particular, the BN trend is

$$\hat{\tau}_t^{BN} = \lim_{h\to\infty} \left\{ E[y_{t+h}|\Omega_t] - h \cdot E[\Delta y_t] \right\}, \tag{4}$$

The BN trend is equivalent to the long-horizon conditional forecast of the level of output minus any deterministic drift. In particular, as the forecasting horizon extends to infinity, a long-horizon forecast of output should no longer be influenced by the realization of its

[5] See, 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.

transitory component at time t. Thus, BN trend should only reflect the expected impact of the permanent component of output 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 (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

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

where

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

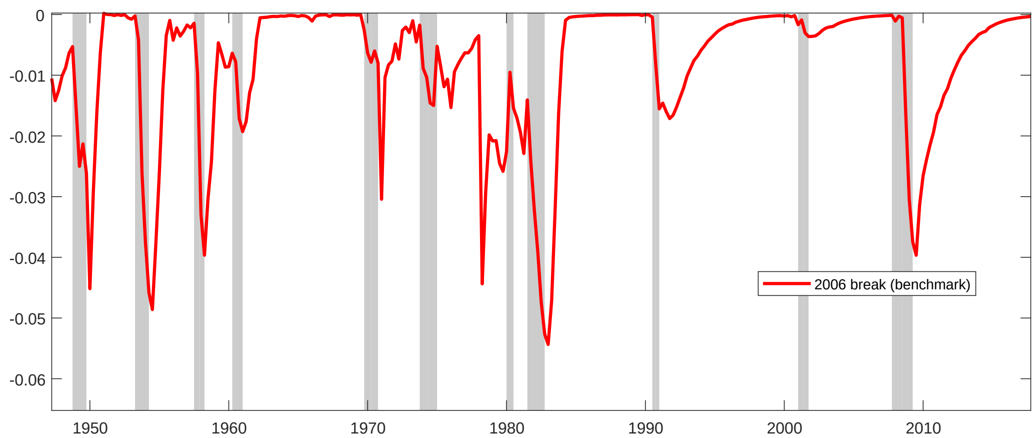
$\tilde{S}_t = (S_t, ..., S_{t-m})'$ is a vector of relevant current and past regimes for forecasting output, $p(\cdot)$ is a probability distribution based on the regime-switching model, S_t is the latent state variable in (1) that evolves according to a fixed transition matrix (2), and i^* is the "normal" 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 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 $p(\tilde{S}_t|\Omega_t)$ 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 calculated as

$$\hat{c}_t^{RDSS} = y_t - \hat{\tau}_t^{RDSS}. \tag{7}$$

Figure 5 plots the estimated output gap for the benchmark specification in (3). We can

[6]See 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.

see that large negative movements in the output gap closely match up with some NBER recessions. However, because an L-shaped recession only affects trend output, the large negative movements in the output gap correspond primarily to the recessions with a high probability of the "plucking" regime in Figure 4.

In terms of the Great Recession, the probability of contraction spikes up and the output gap drops 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. This delayed timing of the severe contraction for the Great Recession is distinct from the behavior 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]

[7] Instead, it may be related to a typical 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.

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 a break in 2006 is supported by the data and plays the crucial role in explaining why the level and growth of U.S. real GDP have been so much lower than would have been projected otherwise 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 Trend Growth

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:Q4 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 (1999*b*) 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 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

[8]The 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: Sequential Structural Break Tests for Output Growth

# of breaks	Estimated Break Dates	LR Test Stat	Critical Value (5%)
1	1984:Q2	68.12	12.09
2	1984:Q2, 2006:Q1	21.93	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	0.37	0.62	[1994:Q4,2007:Q2]
		(b) Restricted Model		
1		0.82	1.16	
2	1984:Q2	0.82	0.49	[1982:Q1,1987:Q2][1991:Q1]
3	2006:Q1	0.37	0.62	[1994:Q4,2007:Q2]

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

1987:Q1, while the confidence set for the second break date is wider and ranges from 1994:Q4 to 2007:Q2.[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.

For the first estimated break in 1984:Q2, a likelihood ratio test of no change in mean suggests 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.

[9]We 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 2007:Q2 represents the last possible break date given 15% trimming and the confidence set would extend to the last possible break date given 10% trimming. However, 2006:Q1 remains the estimated break date even given 10% trimming.

Table 4: Structural Break Tests for Consumption Growth

# of breaks	Estimated Break Dates	LR Test Stat	Critical Value (5%)
1	2000:Q4	46.15	12.09
2	1992:Q1 2006:Q4	28.19	13.39
3	1958:Q1 1992:Q1 2006:Q4	20.32	14.28
4	1958:Q1 1973:Q1 1992:Q1 2006:Q4	13.49	14.87

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 $\hat{\delta} = -0.50$ for the reduction in trend growth from the benchmark specification of our Markov-switching model reported in Table 1.

If there is a break in the trend growth in U.S. real GDP, it should also show up in other time series that share a common trend and balanced growth with aggregate output. Motivated by this, we first consider data for consumption of nondurables and services, as in Eo and Morley (2015).[10] This measure of consumption appears to have a balanced long-run relationship with output corresponding to a cointegrating vector of $(1 -1)$ for the natural logarithms of consumption and output. For consumption growth, we find evidence of three breaks, which are estimated to have occurred in 1958:Q1, 1992:Q1, and 2006:Q4, as reported in Table 4.[11]

Estimates for the mean and standard deviation of consumption growth based on the estimated break dates, along with the confidence sets for the break dates, are reported in Table 5. The first two breaks appear to be in volatility only and their confidence sets are quite wide even for a restricted model that imposes this, ranging from 1957:Q3 to 1981:Q3

[10]The raw data are seasonally adjusted and were obtained from FRED for the same sample period as real GDP. Given additivity not holding for chain-weighted measures, we construct our measure of real consumption of nondurables and services using a Tornqvist approximation that requires real and nominal data for nondurables and for services.

[11]We also consider breaks in a vector error correction model of consumption and output, as in Eo and Morley (2015), but do not report here to conserve space. The multivariate results are similar to a combination of the univariate results for output and consumption, as was also found in Eo and Morley (2015). However, the univariate results have the benefit of a clearer interpretation as breaks in trend growth and/or volatility.

Table 5: Estimates for Mean and Standard Deviation of Consumption Growth Allowing for Structural Breaks

Regime	Estimated Break Date	Mean	Std. Dev.	Confidence Set for Break Date
		(a) Unrestricted Model		
1		0.75	0.78	
2	1958:Q1	0.84	0.50	[1957:Q3,1981:Q3]
3	1992:Q1	0.77	0.27	[68:Q3] [72:Q4,73:Q3] [74:Q1] [75:Q2,75:Q3]
				[76:Q1,79:Q1] [80:Q2,90:Q1] [90:Q4,96:Q1]
4	2006:Q4	0.36	0.35	[2002:Q3,2007:Q2]
		(b) Restricted Model		
1		0.80	0.78	
2	1958:Q1	0.80	0.50	[1957:Q3,1981:Q3]
3	1992:Q1	0.80	0.27	[1980:Q2,1995:Q4]
4	2006:Q4	0.36	0.35	[2002:Q3,2007:Q2]

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

for the first break and 1980:Q2 to 1995:Q4 for the second break. This lack of precision could reflect slower moving changes in consumption volatility over a large portion of the sample. However, the third break clearly corresponds to a more sudden reduction in trend growth and the confidence set is more precise than for the output data, ranging from 2002:Q3 to 2007:Q2. The more precise confidence set for a break in trend growth for consumption than output is intuitive given that the break is estimated to be almost the same magnitude (as would be expected given the balanced long-run relationship with output), with average growth declining by 0.44 in 2006:Q4, while consumption growth volatility is much lower than output growth volatility. Thus, the relative magnitude of the break in terms of its "signal to noise" ratio is much larger for consumption growth than output growth and it should be easier to determine its timing empirically.

One possible concern with the structural break tests for output or consumption is that the regression model framework of Qu and Perron (2007) may not be appropriate given nonlinearity in the underlying data generating processes. To address this, we also consider data for final sales.[12]. The natural logarithms of output and final sales very clearly have a

[12]The raw data are seasonally adjusted and were obtained from FRED for the same sample period as real GDP and aggregate consumption. We consider real final sales.

Table 6: Schwarz Information Criteria

	AR(2)	Markov-Switching Model
Output	739.94	732.72
Consumption	423.97	405.64
Final Sales	653.83	661.64

Note: The SIC is calculated as $-2\ln\hat{L} + k\ln T$, where $\hat{L}$ is the maximized value of the likelihood function and k is the number of parameters. For a given variable, smaller values are preferred. The AR(2) model is for the growth rate of each variable. The Markov-switching model is also for the growth rate and is our benchmark model in (3).

Table 7: Structural Break Tests for Final Sales Growth

# of breaks	Estimated Break Dates	LR Test Stat	Critical Value (5%)
1	1985:Q3	49.050	12.09
2	1985:Q1 2006:Q1	19.364	13.39
3	1958:Q1 1985:Q3 2006:Q1	9.929	14.28

balanced long-run relationship with a cointegrating vector of $(1 -1)$ and almost no difference in their levels on average (their difference, which is related to inventory investment, is close to zero on average). Thus, they should share the same trend growth and we do not need to worry about the effects of possible breaks in the level of their long-relationship, as we might with consumption and output if there happen to be permanent changes in the consumption rate. Furthermore, going back to Sichel (1994), it is widely recognized that apparent non-linearities in output growth are less evident in final sales growth. Consistent with this, the Schwarz Information Criterion (SIC) favors a simple AR(2) model for final sales growth over our Markov-switching model, while it favors the Markov-switching model for both output growth and consumption growth, as reported in Table 6. Thus, the linear regression model framework of Qu and Perron (2007) is arguably more appropriate for final sales growth.

For final sales growth, we find evidence of two breaks, which are estimated to have occurred in 1985:Q3 and 2006:Q1, as reported in Table 7.[13] Meanwhile, estimates for the

[13] We also consider breaks in a vector error correction model of final sales and output, but do not report here to conserve space. As with consumption and output, the multivariate results are similar to a combination of the univariate results for output and final sales, but again the univariate results have the benefit of a clearer interpretation as breaks in trend growth and/or volatility.

Table 8: Estimates for Mean and Standard Deviation of Final Sales Growth Allowing for Structural Breaks

Regime	Estimated Break Date	Mean	Std. Dev.	Confidence Set for Break Date
		(a) Unrestricted Model		
1		0.88	0.94	
2	1985:Q1	0.79	0.48	[1980:Q2,1994:Q2]
3	2006:Q1	0.37	0.51	[1996:Q4,2007:Q2]
		(b) Restricted Model		
1		0.82	0.94	
2	1985:Q1	0.82	0.49	[1980:Q2,1994:Q3]
3	2006:Q1	0.37	0.51	[1997:Q1,2007:Q2]

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

mean and standard deviation of final sales growth based on the estimated break dates, along with the confidence sets for the break dates, are reported in Table 8. The results are very similar to those for output growth, although the confidence set for the break in volatility is less precise. The confidence set for the break in trend growth is somewhat more precise than in the case of output, but less so than for consumption, again consistent with the relative volatilities of the different series.

Taken together, these results provide strong econometric support for a break in trend growth in 2006. Next, we turn to how important accounting for this break is for inferences about the output gap and trend output.

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

When considering different assumptions about structural breaks, we find it is empirically necessary to consider a more general version of our Markov-switching model that allows a partial bounceback effect in the "hysteresis" contractionary regime in addition to the full bounceback effect in the"plucking" contractionary regime. The more general model is given as follows:

Table 9: Maximum Likelihood Estimates under Different Structural Break Assumptions

	2006 Break		1973 Break		No Break	
Parameter	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.
p_{01}	0.0290	(0.0225)	0.0037	(0.0042)	0.0068	(0.0067)
p_{02}	0.0331	(0.0171)	0.0449	(0.0172)	0.0424	(0.0172)
p_{11}	0.7345	(0.1290)	0.9910	(0.0142)	0.9900	(0.0136)
p_{22}	0.8036	(0.0835)	0.6994	(0.1054)	0.6934	(0.1179)
σ^2	0.4328	(0.0499)	0.4701	(0.0462)	0.4887	(0.0481)
μ_0	0.9591	(0.0772)	0.9826	(0.0679)	0.8259	(0.0468)
μ_1	-1.0988	(0.4286)	-2.1004	(0.4743)	-2.6924	(0.4658)
λ_1	-0.0201	(0.1028)	0.3229	(0.0829)	0.4033	(0.0775)
μ_2	-1.9561	(0.1848)	-1.8658	(0.1740)	-1.7741	(0.2279)
λ_2	0.3260	(0.0308)	0.3110	(0.0290)	0.2957	(0.0380)
δ	-0.5030	(0.1312)	-0.2600	(0.0898)		
log-lik	-343.77		-345.05		-348.98	

Note: The 2006 and 1973 break specifications allows for a structural break in 2006:Q1 and 1973:Q1 in (8), respectively. The standard errors of the parameter estimates are reported in parentheses. The sample period is 1947:Q2 to 2017:Q4. We report estimates for both μ_2 and λ_2, although one implies the other given the restriction $\mu_2 + m \cdot \lambda_2 = 0$.

$$\begin{aligned}\Delta y_t = \mu_0 \quad &+ \quad \delta \mathbf{1}(t > T_b) \\ &+ \quad \mu_1 \mathbf{1}(S_t = 1) + \lambda_1 \sum_{k=1}^{m} \mathbf{1}(S_{t-k} = 1) \\ &+ \quad \mu_2 \mathbf{1}(S_t = 2) + \lambda_2 \sum_{k=1}^{m} \mathbf{1}(S_{t-k} = 2) + e_t, \end{aligned} \tag{8}$$

with the presence of λ_1 making this version of the model more general than (8). Unlike λ_2, which is still constrained such that $\mu_2 + m \cdot \lambda_2 = 0$, we leave λ_1 unrestricted in estimation. Also, we consider three cases for the break date T_b. In addition to our benchmark specification of a 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 (i.e., $T_b > T$).

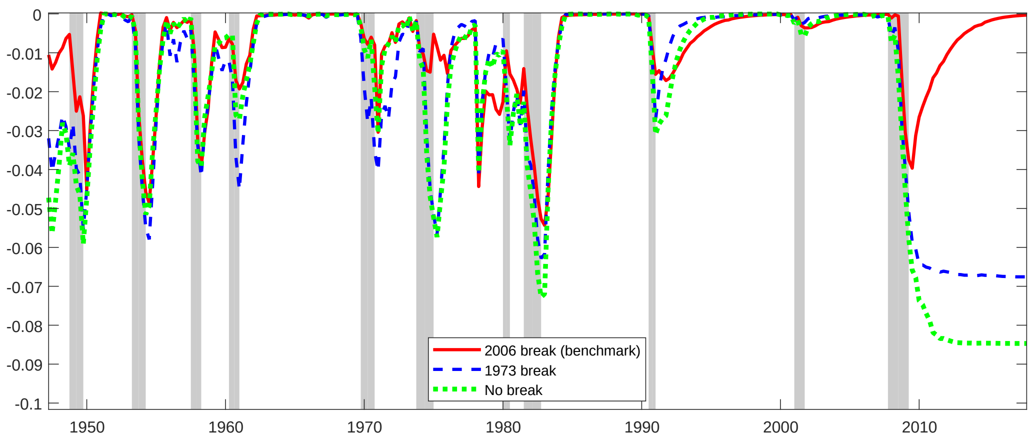
Table 9 reports maximum likelihood estimates of the more general Markov-switching model under different structural break assumptions. The results for the 2006 break case are almost the same as in Table 1, suggesting that our restriction of a strict L shape for the "hysteresis" regime by setting $\lambda_1 = 0$ is warranted in the benchmark specification. In particular, $\hat{\lambda}_1 = -0.02$ and is insignificant, while the estimates of the other parameters are almost identical to before. By contrast, λ_1 is significantly different than zero for both the 1973 break and no break cases, suggesting a partial bounceback effect in the "hysteresis" regime.[14] The other main difference in the estimates for these alternative cases is that the probability of entering or leaving the "hysteresis" regime is much lower than in the 2006 break case. In particular, as we will see, it appears that the "hysteresis" regime in these alternative cases is effectively acting as an absorbing state since the Great Recession, which is consistent with the omission of an important structural break in the model (i.e., the break in 2006 implied by our structural break tests).

Figure 6 plots the estimated output gaps implied by our model under different assumptions about a structural break in trend growth. 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:Q4. To capture this in the 1973 break case and the no break case, the estimated output gap remains persistently large and negative until the end of the sample, while the estimated output gaps are reasonably similar in all three cases prior to 2006. In particular, 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 case and $\hat{\mu}_0 = 0.83$ for no break case. 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 case and 0.30 for no break case, respectively. For both cases, the "hysteresis" 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 a partial bounce-back effect for the "hysteresis" regime, the estimated mean growth rates after 6 quarters for this regime are $\hat{\mu}_0 + \hat{\delta} + \hat{\mu}_1 + 6\hat{\lambda}_1 = 0.55$ (after the break) for the 1973 break case and $\hat{\mu}_0 + \hat{\mu}_1 + 6\hat{\lambda}_1 = 0.55$ for the no break case.[15] These values are consistent

[14] If we impose $\lambda_1 = 0$ in these cases, the model actually identifies fairly similar regimes to the benchmark case, but the fit is considerably worse.

[15] Consistent with its label, the "hysteresis" regime still has a permanent, negative effect on the level of output because $\hat{\mu}_1 + 6\hat{\lambda}_1 < 0$ in both cases.

Figure 6: Estimated Output Gaps for Different Structural Break Assumptions



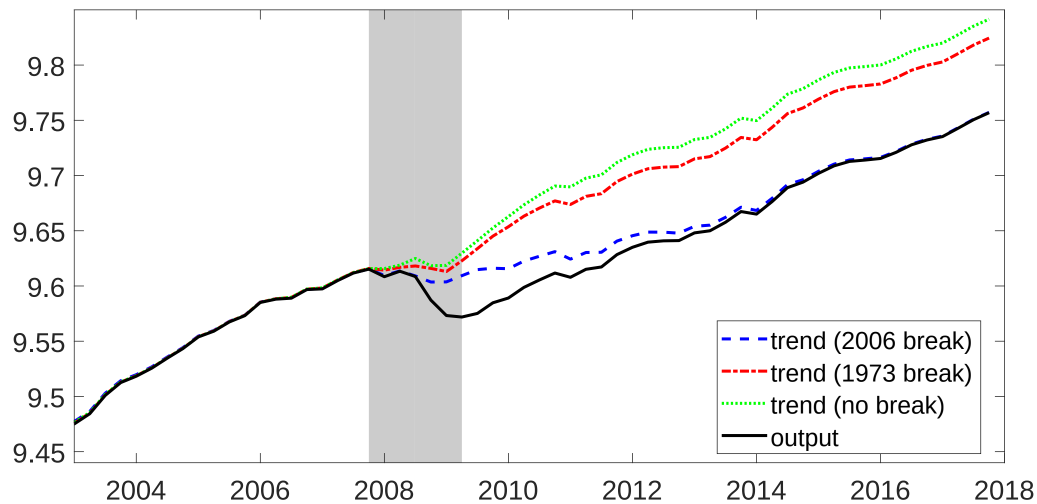
Note: The shaded areas denote NBER recession dates.

with the average growth rate of output since the Great Recession. Notably, however, the likelihood values for these alternative cases 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 the previous subsection 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 cases 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 case and the no break case, 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 low 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.

These results suggest that allowing for a break in trend growth in 2006 is crucial for our inferences about the output gap and trend output since the Great Recession. Although our benchmark specification with a break in 2006 fits better than the alternative cases, we acknowledge that the output gap and trend output could, in reality, be more like the

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



reported estimates when allowing for a break in 1973 or no break. We test this possibility by comparing inflation forecasts for the different cases next.

4.3 Forecasting Inflation

We 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 from the previous subsection. 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

$$\pi_{t+h} - \pi_t = \alpha + \sum_{j=0}^{p-1} \phi_j \Delta\pi_{t-j} + \kappa \hat{c}_t^{RDSS} + \epsilon_{t+h,t}, \tag{9}$$

where π_t is inflation and $\hat{c}_t^{RDSS}$ is the estimated output gap in (7) that depends on the structural break assumption. Inflation is calculated as 100 times the first difference of the natural logarithms of headline and core measures of the Personal Consumption Expenditures

Table 10: Inflation Forecast Evaluation

	Headline PCE Inflation							
	h=1		h=2		h=3		h=4	
	RRMSE	DM	RRMSE	DM	RRMSE	DM	RRMSE	DM
1973 Break	1.50	0.00	1.37	0.01	1.35	0.03	1.62	0.05
No Break	1.67	0.00	1.55	0.00	1.56	0.02	1.96	0.04
	Core PCE Inflation							
	h=1		h=2		h=3		h=4	
	RRMSE	DM	RRMSE	DM	RRMSE	DM	RRMSE	DM
1973 Break	1.76	0.00	2.16	0.00	2.58	0.01	2.55	0.03
No Break	2.01	0.00	2.61	0.00	3.14	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.

(PCE) deflator for the sample period of 1959:Q2 to 2017:Q4.[16] We consider the core PCE inflation rate in addition to the headline PCE inflation rate because the Federal Reserve tends to focus on it as the key measure for tracking underlying inflationary trends.[17] We calculate out-of-sample forecasts over the evaluation sample of 2009:Q3 to 2017:Q4 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 10 show that the estimated output gap for our benchmark specification outperforms the other estimates at all horizons ($h = 1, ..., 4$ 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 proposed in Ball and Mazumder (2011), to capture the behavior of inflation since the Great Recession. However, it is clearly important to have an estimated

[16] The raw data are seasonally adjusted and were obtained from FRED. We consider inflation data from 1959:Q2 only because of the limited availability of the core PCE price deflator.

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

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. 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 fully recovered from the U-shaped recession 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 to be either L or U shaped and fits the U.S. data well, with a clear characterization of most postwar recessions as being 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 behavior of inflation since the Great Recession. However, it is necessary to account for the structural break in trend growth in 2006.

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