THE UNIVERSITY OF SYDNEY

Economics Working Paper Series

2017 - 14

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

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June 2019

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June 21, 2019

Abstract

Since the Great Recession, U.S. real GDP has not returned to its previously projected path, a phenomenon widely associated with secular stagnation. We investigate whether this stagnation is due to hysteresis effects from the recession, a persistent negative output gap following the recession, or slower trend growth for other reasons. To do so, we develop a new Markov-switching time series model of output growth that accommodates two different types of recessions, those which permanently alter the level of real GDP and those with only temporary effects. We also account for structural change in trend growth. Estimates from our model suggest that the Great Recession generated a large persistent negative output gap rather than any substantial hysteresis effects, with the economy eventually recovering to a slower-growth trend path due to an apparent reduction in productivity growth that began sometime prior to the onset of the Great Recession.

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

JEL classification: C22; C51; E32; E37

*We thank James Bullard, Ana Galvao, Nicolas Groshenny, Yu-Fan Huang, Susumu Imai, Ryo Jinnai, Chang-Jin Kim, Ian King, Mengheng Li, Charles Nelson, Masao Ogaki, Jeremy Piger, Ben Wang, and conference and seminar participants at the 2018 Norges Bank Workshop on Nonlinear Models in Macroeconomics and Finance, 2018 SNDE meetings in Tokyo, the 2018 International Symposium on Econometric Theory and Applications in Sydney, the 2018 IAAE conference in Montreal, the 2019 Renmin University of China School of Finance Workshop, the Bank of Japan, University of Adelaide, Hitotsubashi University, Hokkaido University, Keio University, Macquarie University, Universitat Pompeu Fabra, University of Queensland, and University of Technology Sydney for helpful comments and suggestions. This research was supported by ARC DP190100202. The usual disclaimers apply.

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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 different theories of secular stagnation, but Summers (2014, 2015) emphasizes the role of inadequate demand. According to his view, 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 prevented a further lowering of interest rates, inadequate demand caused 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; Guerron-Quintana and Jinnai, 2019) 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 such effects, 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 recessions almost always have negative permanent effects on the level of aggregate 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, 2016), is that it reflects supply-side forces such as slower productivity growth and demographic changes 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 related to slow growth of total factor productivity and a decline in labor force participation, with both phenomena starting before the onset of the recession and not obviously connected to the financial crisis. Supporting this view, a number of recent empirical studies have estimated a structural break in U.S. trend 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 (2018). However, an ability to reject that the slowdown actually occurred during the Great Recession, not before, is unclear from this literature.

In this paper, we develop a highly flexible nonlinear time series model that allows us to examine the empirical support for competing views surrounding why U.S. real GDP has not returned to its projected path prior to the Great Recession. In particular, we investigate whether this stagnation is due to level and growth hysteresis effects from the recession, a persistent negative output gap following the recession, or slower trend growth for other reasons. 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 real GDP growth allows a given recession to either permanently alter the level of aggregate output (an 'L-shaped' recession) or only have a temporary effect (a 'U-shaped' recession).[1] We also account for structural change in trend growth. In particular, using the testing procedure from Qu and Perron (2007), we find an estimated reduction in the long-run growth rate of U.S. real GDP in 2006Q1. 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 rather than a level hysteresis effect, with the economy eventually recovering to a lower-growth trend path. Our finding about the nature of the Great Recession is robust to allowing for more complicated patterns of structural change in trend growth or even assuming no structural change, while the precision of our inference that the break occurred before the Great Recession is sharpened considerably by taking into account nonlinear dynamics. Notably, we are able to formally reject that the slowdown in trend growth occurred after 2006Q2. Furthermore, we find that the apparent timing of the slowdown is more consistent with a reduction in productivity growth than changes in population growth or labor force participation.

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

[1] The univariate approach often taken in the literature on nonlinear output growth dynamics implicitly assumes a common propagation for all underlying symmetric shocks to aggregate output. However, it has the compensating advantage of allowing for a tightly parameterized, but still sophisticated specification of dynamics for asymmetric shocks, in our case different dynamics for two types of recessions.

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, those models make assumptions about the correlations between permanent and transitory movements and implicitly place strong restrictions on the variance of the stochastic trend in aggregate output that do not appear to be supported by the data (Morley, Nelson and Zivot, 2003; Morley, 2007).

The rest of this paper proceeds as follows: In Section 2, we discuss background evidence for nonlinearity and structural breaks in U.S. real GDP. In Section 3, we present the details of our new Markov-switching model and show how it can generate both L- and U-shaped recessions. In Section 4, we report estimates for a benchmark version of our model and examine implications for why real GDP has stagnated since the Great Recession. In Section 5, we consider some extensions of our benchmark model in order to investigate the robustness and interpretation of our results. Section 6 concludes.

2 Background

There is some existing evidence for Markov-switching nonlinear dynamics in U.S. real GDP growth. Specifically, Morley and Piger (2012) formally test for nonlinearity using the procedure developed in Carrasco, Hu and Ploberger (2014) and find support for the Markov-switching model in Kim, Morley and Piger (2005) that captures U-shaped recessions, but not for the model in Hamilton (1989) that captures L-shaped recessions. However, the tests are applied using data over the sample period of 1947-2006 and so do not include the Great Recession. More recently, Morley and Panovska (2019) conduct tests for nonlinearity using data for a number of countries and find similar results to Morley and Piger (2012) of greater support for a Markov-switching model with U-shaped recessions than L-shaped recessions. Notably, for the U.S. data over the sample period of 1947-2016, there is evidence for nonlinearity when allowing for an estimated slowdown in trend growth in 2000Q2 based on Bai and Perron (1998, 2003) testing procedures.[2] As discussed below when presenting our model, these previous studies consider Markov-switching models with only one type of recession, while Eo and Kim (2016) are able to reject these models in favour of more heterogeneity in business cycle regimes. Applying the Carrasco, Hu and Ploberger (2014) testing procedure to our new Markov-switching model with two different types of recessions would not be straightforward, but we are able to show how well the estimated nonlinear dynamics hold up given a very flexible structure and more years of data, including enough observations after the end of the Great Recession to discriminate between competing hypotheses about its long-run consequences.

Before presenting the details of our new Markov-switching model, we follow Morley and Panovska (2019) by first considering possible structural breaks in trend growth in a nonparametric setting. We do so by applying Qu and Perron (2007) testing procedures for multiple structural breaks in mean and/or variance of quarterly U.S. real GDP growth for the sample period of 1947Q2 to 2018Q4 with 10% trimming at the beginning and the end of the sample and between breakdates.[3] The test regression for output growth includes only a constant, but we nonparametrically allow for serial correlation, such as would be implied by our Markov-switching model, in calculating test statistics by employing a heteroskedasticity and autocorrelation consistent estimator of the long-run variance following Andrews and Monahan (1992). Based on a likelihood ratio test, we find evidence of two breaks, which are estimated to have occurred in 1984Q2 and 2006Q1, as reported in Table 1. These breakdates align with the timing of the Great Moderation widely reported in the literature (Kim and Nelson, 1999*b*; McConnell and Perez-Quiros, 2000) and the breakdate for the slowdown in

[2]A minimum length 'trimming' restriction for subsamples of 15% of the total sample when testing for structural breaks means that the estimated breakdate cannot correspond to the Great Recession for the sample period considered in Morley and Panovska (2019), while it can in our analysis given the availability of a few extra years of data, as well as our consideration of 10% trimming. It should also be noted that our reporting their breakdate as 2000Q2 corresponds to the convention of a breakdate being the last period of the previous structural regime. Also, the Bai and Perron (1998, 2003) procedures only allow for a break in mean, but not variance, unlike the Qu and Perron (2007) procedures that we consider in our analysis.

[3]The raw data for seasonally-adjusted quarterly U.S. real GDP were obtained from the St. Louis Fed database (FRED) and quarterly growth rates were calculated as 100 times the first differences of the natural logarithms of the levels data.

Table 1: Sequential structural break tests for output growth

# of Breaks	Test Statistic	5% Critical Value	Estimated Breakdate(s)
1	72.87	12.80	1984Q2
2	18.76	13.96	1984Q2, 2006Q1
3	8.77	14.84	1984Q2, 2000Q2, 2009Q2

Notes: The test regression for output growth includes only a constant, but we nonparametrically allow for serial correlation and heteroskedasticity in the residual. Critical values are from Qu and Perron (2007).

trend growth that was also found in Luo and Startz (2014) and Kamber, Morley and Wong (2018). The structural breaks are significant at the 5% level and there is no support for an additional break at even a 10% level. Related to the Great Moderation and our Markov-switching model, we note that a larger variance for output growth before 1984Q2 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 2018 (34 years). Thus, we will use our Markov-switching model to check whether this estimated structural break is due to the less frequent realization of recessions since 1984 or a reduction in residual volatility.

Estimates for the mean and standard deviation of output growth based on the estimated breakdates, along with the confidence sets for the breakdates, are reported in Table 2. The confidence set for the first breakdate covers a reasonably short interval of 1982Q1 to 1987Q1, while the confidence set for the second breakdate is wider and ranges from 1991Q3 to 2011Q3.[4] The estimated breakdate of 2006Q1 is consistent with the date for the growth slowdown 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 1984Q2, 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%. The average growth rates before and after the first estimated breakdate of 1984Q2 are very close to each other at 0.89 and

[4]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 breakdates, see the original paper. We note that 2011Q3 represents the last possible breakdate given 10% trimming.

Table 2: Mean and standard deviation of output growth given two structural breaks

Subsample	Mean	Std. Dev.	Estimated Breakdate	Confidence Set for Breakdate
1	0.89	1.16		
2	0.80	0.49	1984Q2	[1982Q1,1987Q1]
3	0.41	0.59	2006Q1	[1991Q3,2011Q3]

Note: Confidence sets for the breakdates are based on the inverted likelihood ratio approach in Eo and Morley (2015).

0.80, respectively, by contrast to the average growth rate of 0.41 after the second breakdate of 2006Q1. The decline in average growth since 2006Q1 could be related to the realization of the most severe recession in the postwar period between 2007-2009. Thus, we will also use our Markov-switching model to check whether this estimated structural break is due to the Great Recession or a more sustained decline in trend growth. Related, we will determine whether explicitly accounting for nonlinear dynamics affects the precision of inferences about the timing of structural breaks.

3 Model

We develop a new univariate Markov-switching model of real GDP growth that accommodates two different types of recessions. In particular, the model builds on the Markov-switching models in Hamilton (1989) and Kim, Morley and Piger (2005) that both assume all recessions have the same dynamics by allowing for two distinct 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 temporary effects, corresponding to a restricted version of the model in Kim, Morley and Piger (2005) that is related to Kim and Nelson (1999*a*). The idea of allowing for distinct types of contractionary regimes is strongly 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).

Extending the specification in Kim, Morley and Piger (2005), we assume that output growth, Δy_t, has the following time-varying mean over the business cycle based on three regimes:

$$\Delta y_t = \mu_0 + \mu_1 \cdot \mathbf{1}(S_t = 1) + \mu_2 \cdot \mathbf{1}(S_t = 2) + \lambda_2 \cdot \sum_{k=1}^{m} \mathbf{1}(S_{t-k} = 2) + e_t, \tag{1}$$

where $\mathbf{1}(\cdot)$ is an indicator function, 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, and $S_t = 2$ for the U-shaped contractionary regime according to transition probabilities $Pr[S_t = j|S_{t-1} = i] = p_{ij}$ for $i, j = 0, 1, 2$, and $e_t \sim N(0, \sigma^2)$. For simplicity and following the empirical results in Hamilton (1989), Kim, Morley and Piger (2005), Morley and Piger (2012), Huang, Luo and Startz (2016), Eo and Kim (2016), amongst others, we abstract from linear autoregressive dynamics in the residual from the nonlinear terms by assuming that e_t is serially uncorrelated.[5]

To identify the contractionary regimes as being associated with two different types of recessions, we assume that the economy does not switch directly from one contractionary regime to another without going through an expansionary regime first. This sequencing of regimes is imposed using restrictions on the regime transition probabilities as follows: $p_{12} = 0$ for the L-shaped regime to U-shaped regime transition and $p_{21} = 0$ for the U-shaped regime to L-shaped regime transition. Thus, the overall regime transition matrix is given by

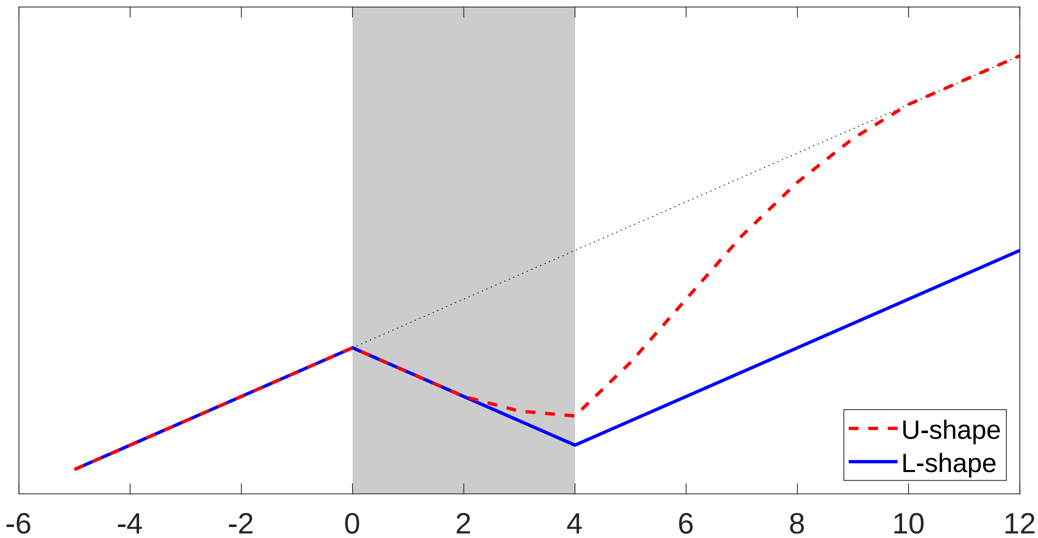
$$\Pi = \begin{bmatrix} 1 - p_{01} - p_{02} & 1 - p_{11} & 1 - p_{22} \\ p_{01} & p_{11} & 0 \\ p_{02} & 0 & p_{22} \end{bmatrix}. \tag{2}$$

The λ_2 parameter is the key distinct feature of the U-shaped contractionary regime in (1) because it allows for a bounceback effect that generates an asymmetric output gap, as in Kim, Morley and Piger (2005); Morley and Piger (2012); Morley and Panovska (2019).[6] To

[5]Specifically, these earlier papers find that linear autoregressive dynamics in the residual are not particularly important once allowing for a Markov-switching mean. Meanwhile, it is important to note that the statistical evidence for Markov-switching nonlinearity discussed above in Section 2 allows for AR(2) dynamics in output growth under the null of linearity. Morley and Rabah (2014) highlight that it is crucial to allow for serial correlation under the null hypothesis in order to avoid any spurious rejections of linearity.

[6]Possible sources of an asymmetric output gap are capacity constraints, monopoly power, asymmetric price adjustments due to collateral requirements, and asymmetric shocks. See Friedman (1964, 1993); DeLong and Summers (1988); Auroba, Bocola and Schorfheide (2013); Guerrieri and Iacoviello (2016); Diebold,

Figure 1: Illustration of different types of recessions



Note: The shaded area denotes the contractionary regime.

clearly identify this regime as distinct from the L-shaped regime, which only has permanent effects on the level of output by construction, we impose the restriction $\mu_2 + m \cdot \lambda_2 = 0$.[7] This restriction implies that, following the realization of $S_t = 2$, the bounceback effect $m \cdot \lambda_2$ exactly cancels out the contractionary effect from μ_2, such that the U-shaped regime only has temporary effects on the level of output, as in the Markov-switching model in Kim and Nelson (1999*a*), but distinct from the model in Kim, Morley and Piger (2005), which does not impose this restriction.[8]

Figure 1 illustrates how the two contractionary regimes create different types of recessions in terms of their long-run effects on the level of output. To demonstrate this, we simulate the path of output implied by the model in (1) before, during, and after the occurrence of a contractionary regime. We set the length of the post-recession bounceback effect to $m = 5$

Schorfheide and Shin (2017), amongst many others, for more information on these theories of business cycle asymmetry.

[7]Typically with Markov-switching models, it is necessary to place labelling restriction such as $\mu_1 < 0$ and $\mu_2 < 0$ to identify the model. However, because there is no bounceback effect when $S_t \neq 2$, the regimes end up being uniquely identified given the restriction on λ_2 and the restrictions on the transition probabilities. Thus, we place no restrictions on the other parameters in (1).

[8]In addition to our consideration of a latent Markov-switching state variable instead of predetermined NBER dates, this restriction on the bounceback effect is another key distinction from Huang, Luo and Startz (2016), who allow for possible permanent effects with their U-shaped regime, as in Kim, Morley and Piger (2005), in addition to assuming permanent effects with their L-shaped regime.

periods and the model parameters to be $\mu_0 = 1$ for the expansionary regime, $\mu_1 = -2$ for the L-shaped regime, and $\mu_2 = -2$ for the U-shaped regime (thus, implying $\lambda_2 = 0.4$ given the restriction we impose to help identify the two different types of recessions). For clarity in seeing the impact of the relative impact of the two different regimes, we abstract from the linear e_t shocks when simulating the path of output. In both cases, we assume that the economy enters a contractionary regime at time $t = 0$ that lasts for 4 quarters and causes a recession. For the U-shaped regime, the bounceback effect takes hold as the recession continues and flattens out the path of output, with the economy then growing quickly and eventually 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 its path traces out what looks like a tilted and elongated "U". By contrast, for the L-shaped regime, the absence of a bounceback 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 its path traces out what looks like a tilted "L".

4 Estimates

Estimation is conducted via maximum likelihood, where the conditional likelihood function given the length of the post-recession bounceback effect m is evaluated based on the filter presented in Hamilton (1989) keeping track of 3^{m+1} states in each period. The estimate of the discrete-value parameter m is also chosen to maximize the likelihood. However, because this is done as a separate step in estimation, it should be noted that the standard errors for the other parameters are only conditional on the estimated value $\hat{m}$.

To incorporate the structural breaks estimated in Section 2 into our benchmark model, we modify the basic model in (1) as follows:

$$\Delta y_t = \mu_0 + \delta \cdot \mathbf{1}(t > \tau) + \mu_1 \cdot \mathbf{1}(S_t = 1) + \mu_2 \cdot \mathbf{1}(S_t = 2) + \lambda_2 \cdot \sum_{k=1}^{m} \mathbf{1}(S_{t-k} = 2) + e_t, \quad (3)$$

where $e_t \sim N(0, \sigma_t^2)$, with $\sigma_t^2 = \sigma_{v0}^2 \cdot \mathbf{1}(t \leq \tau_v) + \sigma_{v1}^2 \cdot \mathbf{1}(t > \tau_v)$. Based on the findings in

Table 3: Estimates for benchmark model

Parameter	Estimate	S.E.
p_{01}	0.026	0.014
p_{02}	0.018	0.010
p_{11}	0.657	0.167
p_{22}	0.728	0.128
μ_0	0.908	0.052
μ_1	-1.320	0.271
μ_2	-2.096	0.288
λ_2	0.419	0.058
δ	-0.407	0.082
σ_{v0}	0.897	0.066
σ_{v1}	0.419	0.027
log-lik	-317.35	

Notes: The benchmark model is given by (3) with structural breaks in trend growth in 2006Q1 and residual volatility in 1984Q2. The standard errors are calculated using numerical second derivatives. The length of the post-recession bounceback effect is estimated to be $m = 5$ and reported standard errors are conditional on this estimate. We jointly estimate μ_2 and λ_2 using the restriction of $\mu_2 + m \cdot \lambda_2 = 0$, but report estimates and standard errors for both parameters.

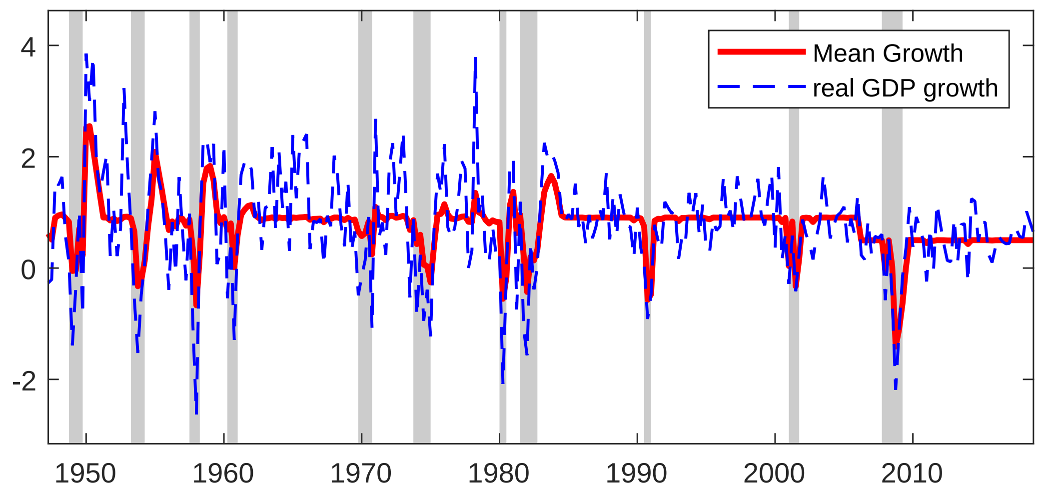
Section 2, the breakdate τ for trend growth and τ_v for volatility in our benchmark model are assumed to be 2006Q1 and 1984Q2, respectively.[9]

4.1 Parameter estimates

Table 3 reports maximum likelihood estimates for the benchmark model. The implied 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 contractionary even though this was not imposed in estimation. The estimated transition probabilities suggest that expansions are much more persistent than both types of recessions, much like the NBER reference cycle. In particular, the implied continuation probability of the expansionary regime $1 - \hat{p}_{01} - \hat{p}_{02}$ is 0.956, with expected duration of 23 quarters, while the expected duration is 3 quarters for the L-shaped regime and 4 quarters for the U-shaped regime. Residual volatility is estimated to have dropped by more than half in 1984Q2, suggesting the Great Moderation was not simply due to a less frequent realization of recessions. Meanwhile, the estimated reduction in trend growth

[9]The breakdate for trend growth is consistent with that assumed in Luo and Startz (2014). We will consider different estimated breakdates and alternative assumptions about structural change in the robustness analysis in Section 5.

Figure 2: Implied time-varying mean and quarterly output growth



Note: The shaded areas denote NBER recession dates.

in 2006Q1 of -0.41 is very close to the reduction of -0.39 found with the Qu and Perron (2007) test in Section 2, suggesting that lower average growth since 2006 was also not simply due to the realization of a severe recession. Meanwhile, the estimated length of the post-recession bounceback effect is 5 quarters, although the other estimates are almost the same for 6 quarters, which was the length considered in Kim, Morley and Piger (2005).[10]

4.2 Time-varying mean and projected output

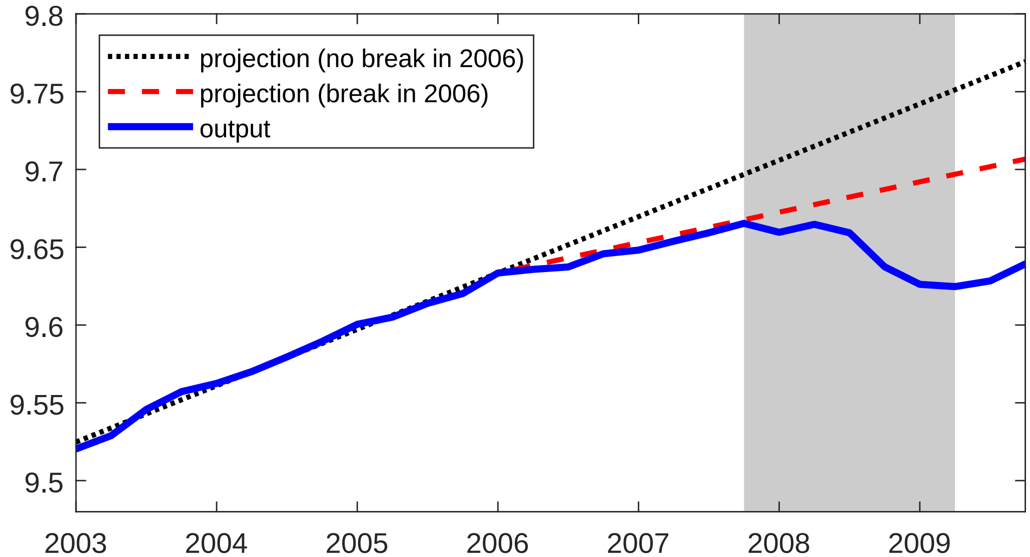
Figure 2 plots the time-varying mean from the benchmark model using the filtered estimate $E[\mu_t|\Omega_t]$, where $\mu_t \equiv \Delta y_t - e_t$ and $\Omega_t \equiv (\Delta y_1, \Delta y_2, ..., \Delta y_t)$. Closely tracking realized real GDP growth and reflecting $\hat{\delta} = -0.41$, the time-varying mean declines abruptly after 2006Q1, with this slowdown in trend growth clearly contributing to the weak recovery of the U.S. economy following the Great Recession.[11]

To illustrate the magnitude of the trend break in 2006Q1, Figure 3 plots projections from $t = $ 2006Q1 for future log output $E[y_{t+h}|\Omega_t]$, $h > 0$, both accounting for and not accounting

[10]For comparison, the log-likelihood values for $m = 4, 6, 7$ are -318.59, -317.65, and -318.68, respectively.

[11]Figure 2 looks similar to 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 also allowing for possible structural change in trend growth. Our simpler model captures differences in mean growth for each recession and expansion based on whether the contractionary regime is L or U shaped, with the mean growth in a recession related to the mean growth in the subsequent expansion.

Figure 3: Projected and realized output



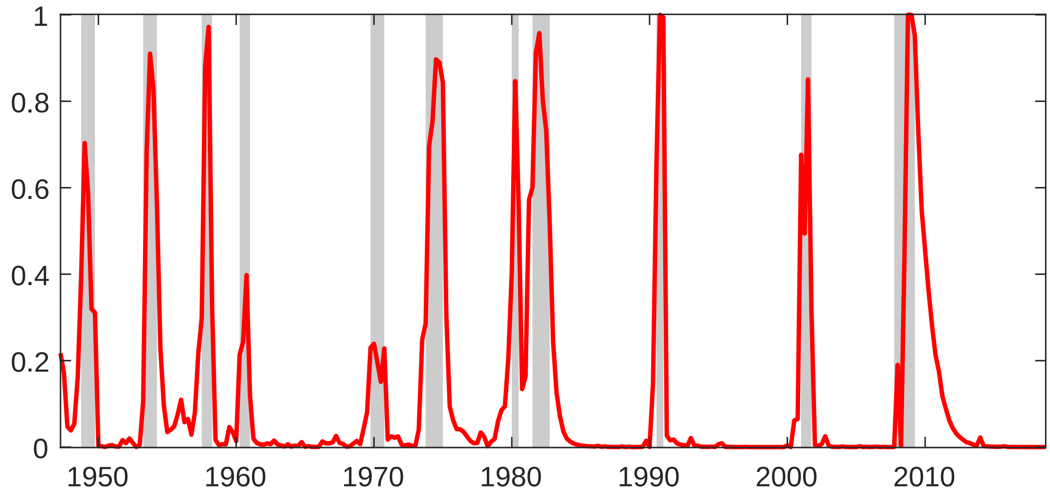
Notes: Output and projections are reported in natural logs. We calculate projections in 2006Q1 (the structural breakdate in output growth) (i) assuming no break with the dotted black line and (ii) accounting for the structural break in output growth with the dashed red line. The shaded area denotes the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

for the structural break. The black dotted line shows the projection of log output 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 decline in trend growth began in 2006 prior to the onset of the Great Recession. Notably, the difference by the end of the Great Recession corresponds to more than 5% of the level of real GDP in 2006Q1.

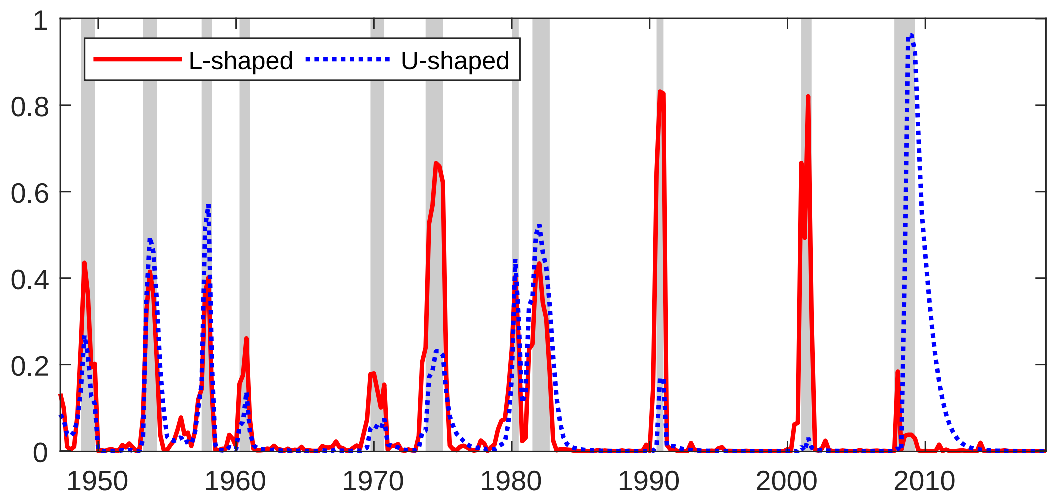
4.3 Classifying recessions as L shaped or U shaped

Figure 4 reports the smoothed probabilities of being in a contractionary regime in period t. The top panel plots the probability of being in one or the other regime, calculated from the sum of the probabilities of being in the L-shaped regime and the U-shaped regime using $Pr[t = \text{contraction}|\Omega_T] \equiv Pr[S_t = 1|\Omega_T] + Pr[S_t = 2|\Omega_T]$. This probability closely matches the timing of NBER recessions. In particular, for nine of the eleven NBER recessions in the sample, the smoothed probability is well above 50% over most of a given recession.

Figure 4: 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. Smoothed probabilities are calculated. The shaded areas denote NBER recession dates.

The bottom panel of Figure 4 plots the underlying smoothed probabilities of the L-shaped and U-shaped regimes. Considering their relative contribution to the overall probability of a contractionary regime, these probabilities suggest that the 1973-75, 1990-91, and 2001 recessions in particular can clearly be classified as L shaped, while the 2007-09 recession can clearly be classified as U shaped. The less definitive classification of the other recessions suggests they may exhibit more of a partial recovery, as found for the estimated bounceback model in Kim, Morley and Piger (2005).

4.4 Permanent effects of recessions on the level of output

Taking the regime probabilities as indicative of the nature of a given recession, we can calculate an implied permanent effect of the j^{th} NBER-dated recession on the level of output as

$$\sum_{t=NBER_{j,s}-1}^{NBER_{j,e}+1} \mu_1 \cdot Pr[S_t = 1|\Omega_T] \tag{4}$$

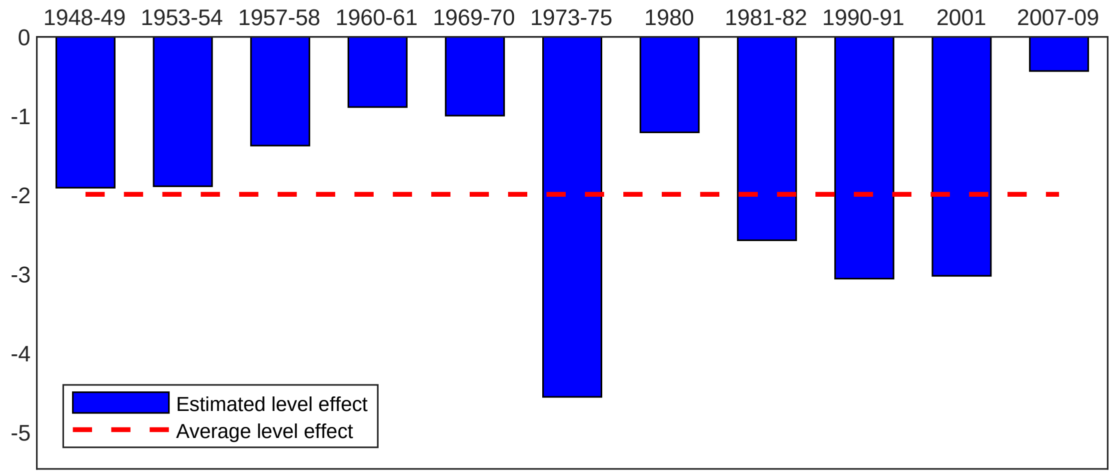
where $NBER_{j,s}$ is the start date of the j^{th} NBER-dated recession and $NBER_{j,e}$ is the end date of the j^{th} NBER-dated recession. Figure 5 reports the estimated permanent effect of each recession on the level of output. Despite its relative length, the Great Recession stands out as having a much smaller level hysteresis effect than the other recessions. Meanwhile, some of the harder to classify recessions have closer to the average effect, while the L-shaped recessions have the largest level hysteresis effects. As with the time-varying mean in Figure 2, these results are consistent with the general finding in Eo and Kim (2016) that not all recessions are alike.

4.5 The output gap

It might perhaps be surprising that the Great Recession is classified as being U shaped given the conventional view that recessions associated with financial crises can have large permanent effects on the level of economic activity.[12] Also, mean growth in Figure 2 does not display the same surge after the Great Recession as occurred after other recessions with a sizeable probability of being U shaped. The explanation for this is that the probability corresponding to a U-shaped regime in Figure 4 remains elevated for a while after the trough date established by the NBER for the Great Recession. This could be related to a prolonged weak labor market ('jobless recovery') following the recession. Also, the zero-lower-bound on interest rates restricted the ability of monetary policy to help stimulate a strong recovery

[12]See, for example, Cerra and Saxena (2008), Reinhart and Rogoff (2009), and Jordà, Schularick and Taylor (2017), amongst many others.

Figure 5: Implied level hysteresis effects of recessions



Notes: Units are % of initial level of output at the onset of the recession.

immediately after the recession. Thus, the relatively tame mean growth following the Great Recession could be related to a large persistent negative output gap that only dissipates very slowly.

To estimate the output gap implied by our model, we consider the Beveridge and Nelson (1981) (BN) decomposition. In particular, the BN trend is

$$\hat{\tau}_t^{BN} = \lim_{h\to\infty} \left\{E[y_{t+h}|\Omega_t] - h \cdot E[\Delta y_t]\right\}. \tag{5}$$

The BN trend is equivalent to the long-horizon conditional forecast of the level of output minus any deterministic drift. As the forecasting horizon extends to infinity, the long-horizon forecast should no longer be influenced by the realization of the output gap at time t. Thus, $\hat{\tau}_t^{BN}$ should only reflect the expected impact of the trend 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

$$\hat{\tau}_t^{RDSS} = \sum_{\tilde{S}_t} \left\{ \hat{\tau}_t^{RDSS}(\tilde{S}_t) \cdot Pr[\tilde{S}_t | \Omega_t] \right\}, \tag{6}$$

where

$$\hat{\tau}_t^{RDSS}(\tilde{S}_t) = \lim_{h \to \infty} \left\{ E[y_{t+h} | \left\{ S_{t+k} = i^* \right\}_{k=1}^{h}, \tilde{S}_t, \Omega_t] - h \cdot E[\Delta y_t | \left\{ S_{t+k} = i^* \right\}_{-\infty}^{\infty}] \right\}, \tag{7}$$

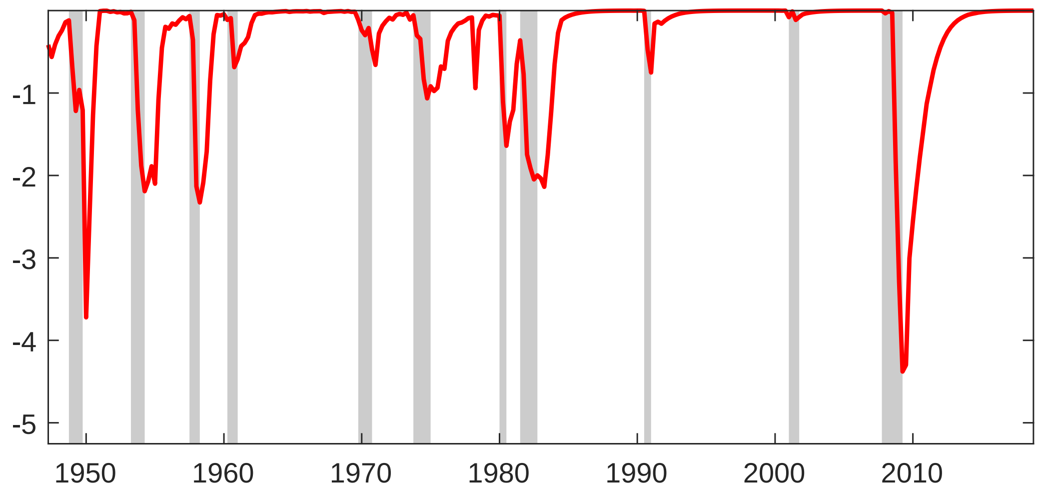
$\tilde{S}_t = (S_t, ..., S_{t-m})'$ is a vector of relevant current and past regimes for forecasting output, $Pr[\cdot]$ is a probability distribution based on the regime-switching model, S_t is the latent state variable in (1) that evolves according to a fixed transition matrix in (2), and i^* is a regime in which the mean of the transitory component is assumed to be zero. Unlike the traditional BN decomposition, there is no implicit assumption that the cycle is unconditionally mean zero and we choose the expansionary regime as having a mean-zero transitory component (i.e., $i^* = 0$).[13] Meanwhile, the probability $Pr[\tilde{S}_t | \Omega_t]$ can be evaluated via the Hamilton (1989) filter. Given $\hat{\tau}_t^{RDSS}$ in (6), the estimated output gap, $\hat{c}_t^{RDSS}$, can be calculated as

$$\hat{c}_t^{RDSS} = y_t - \hat{\tau}_t^{RDSS}. \tag{8}$$

Figure 6 plots the estimated output gap implied by the benchmark model. The large negative movements in the output gap closely match up with some of the NBER-dated recessions. However, because an L-shaped contractionary regime is assumed to only affect trend, the large negative movements in the output gap correspond primarily to the recessions with a high probability of being U shaped. In terms of the Great Recession, the negative output gap opens up later than the NBER peak date of 2007Q4, corresponding to when the probability of U-shaped regime spikes up in Figures 4. As Figure 3 makes clear, the reason for this different timing 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

[13]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 having a mean-zero transitory component.

Figure 6: Estimated output gap



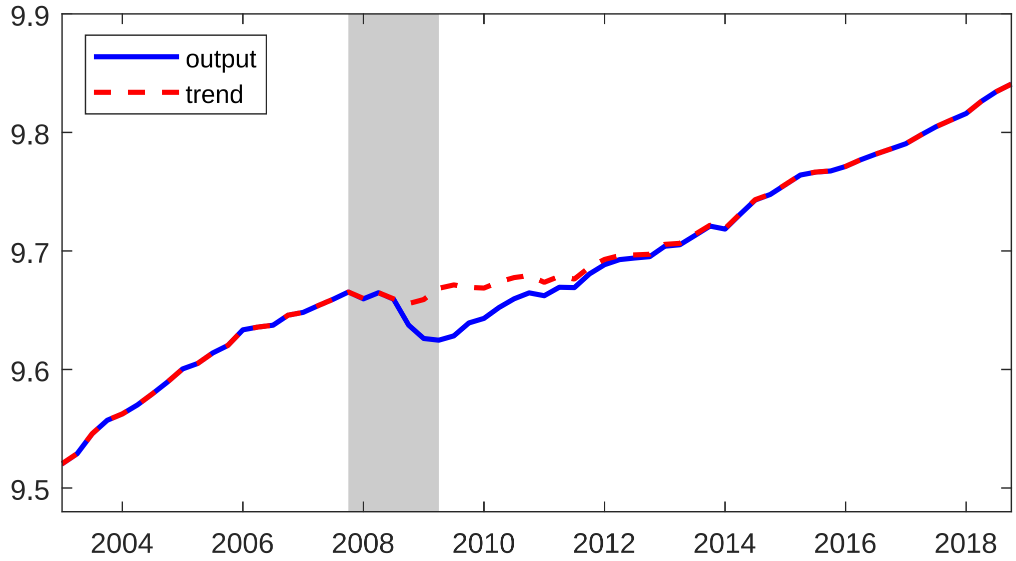
Note: The shaded areas denote NBER recession dates.

timing of the severe contraction for the Great Recession is distinct from the behavior of real GDP in previous recessions and could possibly reflects a misattribution by the NBER of a particularly lackluster manifestation of weak trend growth during the first half of 2008 as being part of the recession phase.[14] Meanwhile, the output gap remains persistently negative long after the NBER trough date, corresponding to only a very slow recovery in the level of output and explaining why there is no surge in growth immediately after the recession in Figure 2.

Figure 7 plots log output and estimated trend from the RDSS decomposition around the Great Recession. The magnitude and persistence of the output gap following the recession is clear from this figure. In particular, the implied negative output gap is not estimated to fully close until around 2012. Because the closure of the output gap is so slow, there is no apparent surge in output growth following the recession in Figure 2. However, it is important to note that this estimated dynamic of a persistent negative output gap is clearly distinctly identified from an L-shaped recession that only alters the level of trend output. If we consider a modification of our model to impose that the Great Recession was L shaped and not U shaped, as found in Huang, Luo and Startz (2016) using NBER dates for the

[14]Instead, the weak growth may be related to a typical end-of-expansion overhiring phenomenon (Gordon, 2003) that could have lowered productivity before the onset of an actual recession in the second half of 2008.

Figure 7: Output and trend around the Great Recession



Note: The shaded areas denote NBER recession dates.

regimes, the fit noticeably deteriorates, with the log-likelihood dropping to -319.61 from -317.35 for our benchmark model.[15] The deterioration of fit appears to be due to a failure to capture the rounded U shape of the recession as it approaches its trough and an eventual gradual recovery of output to a trend path that are both evident in Figure 7.

5 Robustness and Interpretation

In this section, we consider some extensions to our benchmark model in order to investigate the robustness and interpretation of our results. First, we estimate two alternatives to our model that allow us to test whether there really are different types of recessions. Second, we estimate our model using output per capita and examine the role of demographics in driving our results. Third, we estimate breakdates for the structural breaks in trend growth and residual volatility as additional parameters in our model rather than assuming the estimated breakdates from Section 2. Fourth, we check whether our inferences about the

[15] To estimate the modified model that imposes the Great Recession is a contractionary L-shaped regime only, we alter the parameters for the expansionary and U-shaped contractionary regime to take on implausible values for the duration of the NBER dates corresponding to the Great Recession.

Great Recession are robust to alternative assumptions about structural change in trend growth.

5.1 Are there really two different types of recessions?

To test whether there are actually different types of recessions, we consider two alternative models. The first model is more general than our benchmark model in that it allows for a possible bounceback effect in the first contractionary regime in addition to the full recovery in the second contractionary regime:

$$\begin{aligned}\Delta y_t = \mu_0 + \delta \cdot \mathbf{1}(t > \tau) \quad &+ \quad \mu_1 \cdot \mathbf{1}(S_t = 1) + \lambda_1 \cdot \sum_{k=1}^{m} \mathbf{1}(S_{t-k} = 1) \\ &+ \quad \mu_2 \cdot \mathbf{1}(S_t = 2) + \lambda_2 \cdot \sum_{k=1}^{m} \mathbf{1}(S_{t-k} = 2) + e_t, \end{aligned} \tag{9}$$

where the possibility that $\lambda_1 \neq 0$ makes the model more general than in (3). Unlike λ_2, which is constrained such that $\mu_2 + m \cdot \lambda_2 = 0$, we leave λ_1 unrestricted in estimation. Thus, the general model nests our benchmark model if $\hat{\lambda}_1 = 0$. In principle, it also nests the possibility that there are only U-shaped recessions with full recoveries if $\hat{\mu}_1 = \hat{\mu}_2$ and $\hat{\lambda}_1 = \hat{\lambda}_2$, although the regime transition probabilities would not be well identified in this case. The second model is a restricted version of the general model in (9) with only one contractionary regime and corresponds to the original bounceback model in Kim, Morley and Piger (2005):

$$\begin{aligned}\Delta y_t = \mu_0 \quad &+ \quad \delta \cdot \mathbf{1}(t > \tau) \\ &+ \quad \mu_1 \cdot \mathbf{1}(S_t = 1) + \lambda_1 \cdot \sum_{k=1}^{m} \mathbf{1}(S_{t-k} = 1) + e_t. \end{aligned} \tag{10}$$

where, again, we leave λ_1 unrestricted in estimation and we only need to estimate regime transition parameters p_{01} and p_{11}. This restricted model nests the possibility that there are only L-shaped recessions if $\hat{\lambda}_1 = 0$. For both alternative models, $e_t \sim N(0, \sigma_t^2)$ is specified as in (3) to allow for a structural break in residual volatility. The breakdates are the same as in the benchmark model: $\tau = 2006Q1$ and $\tau_v = 1984Q2$. For direct comparability to our benchmark model, we also set $m = 5$ in estimation.

Table 4: Estimates of parameters for alternative models

	General Model		Restricted Model	
Parameter	Estimate	S.E.	Estimate	S.E.
p_{01}	0.028	0.015	0.045	0.017
p_{02}	0.019	0.010		
p_{11}	0.684	0.143	0.801	0.070
p_{22}	0.726	0.127		
μ_0	0.944	0.048	0.940	0.053
μ_1	-1.019	0.223	-1.238	0.146
μ_2	-2.075	0.278		
λ_1	-0.103	0.046	0.121	0.039
λ_2	0.415	0.056		
δ	-0.394	0.079	-0.481	0.077
σ_{v0}	0.902	0.065	0.995	0.073
σ_{v1}	0.409	0.026	0.427	0.027
log-lik	-315.74		-323.87	

Notes: The general model is given by (9) and the restricted model is given by (10). Both models allow for structural breaks in trend growth in 2006Q1 and residual volatility in 1984Q2. For comparability of results to the benchmark model, the length of the post-recession bounceback effect is set to $m = 5$. The standard errors are calculated using numerical second derivatives. Both models allow for a bounceback effect when $S_t = 1$ with $\lambda_1 \neq 0$. For the general model, we jointly estimate μ_2 and λ_2 using the restriction of $\mu_2 + m \cdot \lambda_2 = 0$, but report estimates and standard errors for both parameters.. The restricted model is the same as that proposed in Kim, Morley and Piger (2005) and assumes there are only two regimes.

Table 4 reports maximum likelihood estimates for the two alternative models in (9) and (10). For the general model, the estimate for the additional parameter $\hat{\lambda}_1 < 0$, implying prolonged slow growth following an L-shaped recession rather than a bounceback effect. Given an offsetting smaller magnitude for $\hat{\mu}_1$ and other parameters similar to those in Table 3, the implied dynamic effects of the two types of recessions are close to those in the benchmark model. Furthermore, if we constrain $\lambda_1 \geq 0$, the maximum likelihood estimate is exactly $\hat{\lambda}_1 = 0$, with the other estimates naturally being the same as in Table 3. Meanwhile, for the restricted model, the estimates imply a partial recovery for all recessions given $\hat{\lambda}_1 < -\hat{\mu}_1/5$, which represents an averaging of the effects of the two contractionary regimes for the general model. The fit of the restricted model is considerably worse, although a likelihood ratio test of the two models would not have a standard distribution. These results, combined with the different smoothed probabilities for the two regimes in Figure 4, support the idea of two different types of recessions in the U.S. economy.

Table 5: Estimates of parameters using output growth per capita

Parameter	Estimate	S.E.
p_{01}	0.030	0.014
p_{02}	0.024	0.011
p_{11}	0.751	0.093
p_{22}	0.667	0.131
μ_0	0.644	0.044
μ_1	-1.287	0.134
μ_2	-2.086	0.222
λ_2	0.415	0.056
δ	-0.400	0.073
σ_{v0}	0.869	0.064
σ_{v1}	0.404	0.025
log-lik	-313.69	

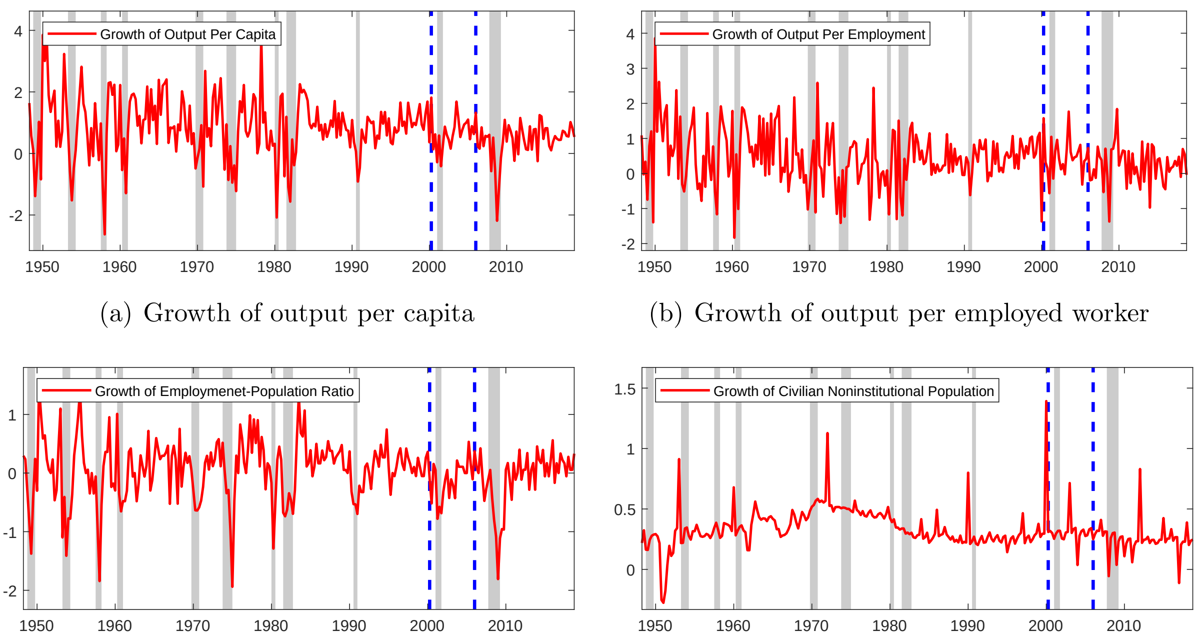
Notes: The model for output growth per capita is the same as the benchmark case in (3) with structural breaks in trend growth in 2006Q1 and residual volatility in 1984Q2. For comparability of results to the benchmark model, the length of the post-recession bounceback effect is set to $m = 5$. The standard errors are calculated using numerical second derivatives. We jointly estimate μ_2 and λ_2 using the restriction of $\mu_2 + m \cdot \lambda_2 = 0$, but report estimates and standard errors for both parameters..

5.2 What role did demographics play in the trend growth slowdown?

We apply our model with two different types of recessions to output growth per capita in order to isolate the effects of population growth on overall trend growth. Table 5 reports the estimates for this case. The estimates are strikingly similar to those for output growth presented in Table 3. One particularly notable similarity is that the slowdown in trend growth per capita is estimated to be $\hat{\delta} = -0.40$, which is very close to $\hat{\delta} = -0.41$ from the benchmark model. This directly implies that population growth is not responsible for the slowdown in overall trend growth since 2006, but instead suggests possible roles for productivity and labor force participation, as argued by Fernald et al. (2017).

Figure 8 plots growth rates of intensive measures and extensive sources of output. In particular, the top two panels correspond to intensive measures and show the growth rates of output per capita and per employed worker, while the bottom two panels correspond to extensive sources and show the population growth rate and the growth of the employment-population ratio. It is clear that there is an abrupt fall in the average level of intensive

Figure 8: Growth rates of intensive output and extensive factors



(a) Growth of output per capita

(b) Growth of output per employed worker

(c) Growth of the employment-population ratio

(d) Growth of population

Notes: The blue dotted lines indicate the dates of 2000Q2 and 2006Q1. The shaded areas denote NBER recession dates.

output in 2006, while there is no obvious fall in the extensive sources of growth at the same time. Population growth does look somewhat lower on average from the onset of the Great Recession and growth of the employment-output ratio falls dramatically during the Great Recession, as it has done in other recessions. However, these changes in population growth and growth of the employment-output ratio cannot explain the estimated timing of a structural break in 2006 and employment-output ratio growth seems to stabilize after the Great Recession. Thus, it appears that demographic factors such as population growth and labour force participation (at least assuming the unemployment rate is stationary) are not responsible for a slowdown in output growth since 2006. Instead, given the slowdown is evident in intensive measures of output growth, it would seem that changes in productivity growth are the most likely source of the slow trend growth in output.

Table 6 reports trend growth decompositions based on basic accounting relationships between the growth rates of output, output per capita, output per employed worker, the employment-population ratio, and population before and after a breakdate in 2006Q1. We

Table 6: Trend growth decompositions

		$\Delta \ln Y_t$	$\Delta \ln(Y_t/N_t)$	$\Delta \ln(Y_t/E_t)$	$\Delta \ln(E_t/N_t)$	$\Delta \ln N_t$
Model trend growth	pre-2006	0.908	0.644	0.452	0.129	
	post-2006	0.501	0.244	0.232	0.034	
	Reduction	-0.407	-0.400	-0.220	-0.095	
Average growth	pre-2006	0.858	0.515	0.468	0.047	0.343
	post-2006	0.405	0.155	0.230	-0.076	0.250
	Reduction	-0.453	-0.360	-0.238	-0.123	-0.093

Notes: The breakdate for the subsamples is 2006Q1. Y_t, N_t, E_t denote output, population, and employment, respectively. Model trend growth corresponds to estimated growth in the expansionary regime in (3) and, given differences in estimated regimes, they do not necessarily add up exactly according to the motivating accounting relationship between the variables given in (11) and (12). Average growth rates do add up exactly.

consider the estimated growth in an expansionary regime implied by our model applied to the various growth rate series, as well as subsample averages.[16] The accounting relationships that inform our trend growth decompositions are

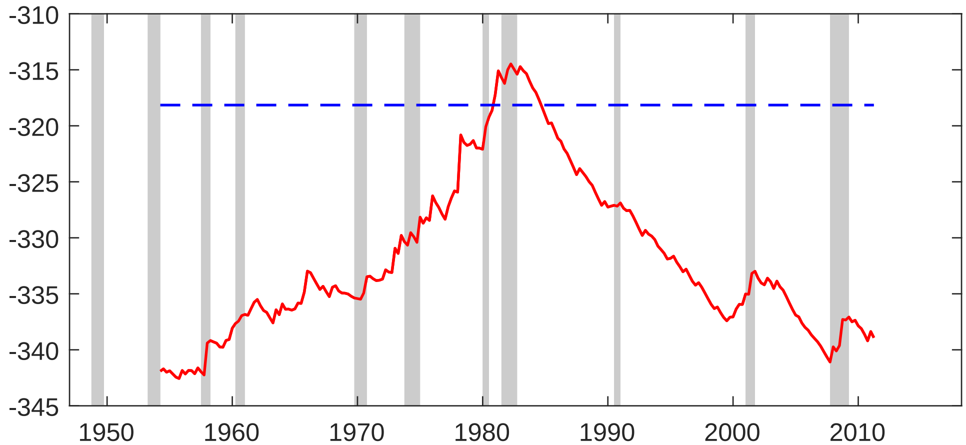
$$\Delta \ln Y_t \equiv \Delta \ln(Y_t/N_t) + \Delta \ln N_t, \tag{11}$$

$$\Delta \ln Y_t \equiv \Delta \ln(Y_t/E_t) + \Delta \ln(E_t/N_t) + \Delta \ln N_t, \tag{12}$$

where Y_t, N_t, E_t denote output, population, and employment, respectively. Corresponding to the results reported in Tables 3 and 5 and the visual impressions in Figure 8, a lot of the slowdown in overall trend growth can be explained by a reduction in the growth rate of output per capita rather than population growth. Indeed, in terms of estimates for our benchmark model, almost all of the slowdown is accounted for by a reduction in trend growth for output per capita. In terms of subsample estimates, most of the slowdown is accounted for in the same way, although we note that the Great Recession has considerable influence on average growth rates since 2006 that is controlled for in our model-based estimates of trend growth. In terms of output per capita growth, more of the slowdown can be explained by a reduction in the growth of output per employed worker than by a reduction in the growth of the employment-population ratio. Thus, these results confirm the impression in Figure 8 that productivity played a bigger role than demographics in explaining the slowdown in

[16]Parameter estimates for our benchmark model applied to the growth rates of output per employed worker and the employment-population ratio are not reported, but are available from the authors upon request.

Figure 9: Profile likelihood for residual volatility breakdate τ_v



Notes: This figure reports the natural logarithm of the profile likelihood function for the volatility breakdate conditional on a trend growth break in 2006Q1. The dashed blue line represents the cutoff based on the 5% critical value for a likelihood ratio test of the breakdate from Eo and Morley (2015). Whenever the profile likelihood value is above the cutoff, the corresponding date is included in the 95% confidence set for the volatility breakdate.

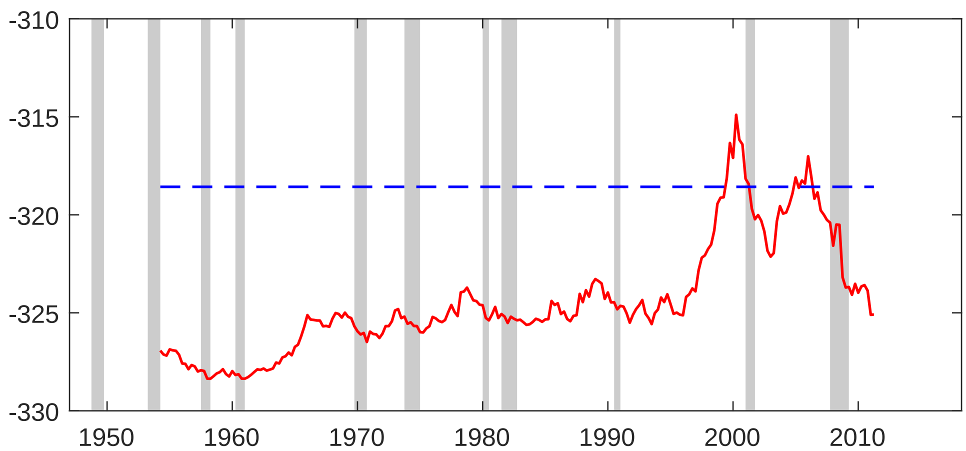
overall trend growth.

5.3 What does our model imply about the timing of structural breaks?

In Section 2, we found breakdates of $\tau = 2006Q1$ and $\tau_v = 1984Q2$ for trend growth and volatility, respectively, using Qu and Perron (2007) testing procedures. Based on this result, we assumed these breakdates in our benchmark model. Here, we examine whether inferences about structural breaks are robust to estimating their timing assuming our Markov-switching model captures the dynamics of output growth.

Figure 9 plots the profile likelihood for the residual volatility breakdate τ_v based on the Markov-switching model in (9) and conditional on a trend growth breakdate of $\tau = 2006Q1$. The structural break in residual volatility is estimated to have occurred in 1982Q2, which is close to the breakdate of 1984Q2 assumed in the benchmark model. The log likelihood value for the volatility breakdate of 1982Q2 is -315.28 compared to the value of -317.35 for the benchmark model with the breakdate in 1984Q2. However, the difference is less than the cutoff value used for constructing a 95% confidence set for a breakdate in Eo

Figure 10: Profile likelihood for trend growth breakdate τ



Note: This figure reports the natural logarithm of the profile likelihood function for trend growth breakdate conditional on a residual volatility break in 1984Q2. The dashed blue line represents the cutoff based on the critical value for a likelihood ratio test of the breakdate from Eo and Morley (2015). Whenever the profile likelihood value is above the cutoff, the corresponding date is included in the confidence set for the trend growth breakdate.

and Morley (2015). Therefore, the confidence set for the volatility breakdate includes the benchmark assumption of 1984Q2 obtained from Qu and Perron (2007) procedures in Section 2. Furthermore, conditional on a break in 1982Q2, we find no support for an additional structural break in residual volatility.

Figure 10 plots the profile likelihood for the trend growth breakdate τ based on the Markov-switching model in (9) and conditional on a residual volatility breakdate of $\tau_v = 1984Q2$. The structural break in trend growth is estimated to have occurred in 2000Q2, which is the same as found in Morley and Panovska (2019) using Bai and Perron (1998, 2003) testing procedures for a shorter sample period and 15% trimming, but earlier than what was found using the Qu and Perron (2007) procedures and 10% trimming in Section 2. Notably, however, our assumed trend growth breakdate of 2006Q1 in the benchmark model is a local mode for the profile likelihood and cannot be rejected using the cutoff value for constructing a 95% confidence set for a breakdate in Eo and Morley (2015). Furthermore, the last date in the 95% confidence set is 2006Q2 and we find no support for an additional structural break in trend growth. Thus, compared to the results for the Qu and Perron (2007) procedures, our Markov-switching model sharpens inferences about the timing of a structural break in trend growth and allows us to clearly reject that the trend growth slowdown occurred with the onset of the Great Recession.

It may be that the structural break in trend growth actually occurred in 2000, as implied by highest mode in Figure 10, making it even clearer that it was not related to the Great Recession. However, it is also possible that the estimated breakdate of 2000Q2 reflects an overfitting of temporary slow growth associated with the 2001 recession, as is evident in Figures 2 and 8. We investigate this further in the next subsection.

5.4 Are inferences about the Great Recession robust to alternative assumptions about structural change?

To the extent that there is uncertainty about the timing of a structural break in trend growth or whether it can even be characterized by a single abrupt change, it is important to investigate the robustness of our inferences regarding the nature of the Great Recession to different assumptions about structural change. To do so, we consider four alternative assumptions: no break; a break in 2000Q2; gradual change addressed by dynamically demeaning output growth rate using a backward-looking rolling 40-quarter average growth rate, as in Kamber, Morley and Wong (2018); and gradual change addressed by using weighted-average inferences based on the relative profile likelihood value over all of the possible breakdates.

Dynamic demeaning involves calculating deviations from a slowly-moving time-varying unconditional mean as follows:

$$\Delta\tilde{y}_t \equiv \Delta y_t - \frac{1}{40}\sum_{j=0}^{39}\Delta y_{t-j}. \tag{13}$$

We then estimate our Markov-switching model using the dynamically-demeaned data $\Delta\tilde{y}_t$ and setting $\delta = 0$ in (3) with the residual volatility breakdate fixed at 1984Q2. Meanwhile, weighted-average inferences involve calculating probabilistic weights over different possible breakdates. In particular, using the relative profile likelihood value for each breakdate, the probabilistic weight for a breakdate τ is calculated as follows:

$$\hat{w}(\tau) \equiv \frac{f(y|\hat{\theta}_\tau;\tau)}{\sum_{t\in[0.1T,0.9T]} f(y|\hat{\theta}_t;t)}, \tag{14}$$

where $f(y|\hat{\theta}_\tau;\tau)$ is the likelihood value for the trend growth breakdate τ given the model in (3) with maximum likelihood estimates $\hat{\theta}_\tau$ for the other parameters conditional on τ and the volatility breakdate again fixed at 1984Q2. By construction, the sum of the weights over the possible breakdates will equal one, $\sum_\tau \hat{w}(\tau) = 1$. Given these weights, the weighted-average smoothed probability of the regime j at time t is given by

$$\sum_\tau \hat{w}(\tau) \cdot Pr[S_t = j|\Omega_T, \tau] \tag{15}$$

and the weighted-average level effect for the j^{th} recession episode is given by

$$\sum_\tau \hat{w}(\tau) \cdot \left\{ \sum_{t=NBER_{j,s}-1}^{NBER_{j,e}+1} \hat{\mu}_{1,\tau} \cdot Pr[S_t = j|\Omega_T, \tau] \right\}, \tag{16}$$

where $Pr[S_t = j|\Omega_T, \tau]$ is the smoothed probability of the regime j at time t given the breakdate of τ. These inferences inherently lose some precision compared to knowing the exact breakdate, but they are potentially robust to multiple or gradual breaks in trend growth.

Table 7 reports the parameter estimates for our Markov-switching model under the following different structural break assumptions: no break; a break in 2000Q2; and gradual change addressed by dynamic demeaning. The parameter estimates related to the effects of recessions are highly robust to the different assumptions about structural change and similar to the estimates for the benchmark model. Thus, our findings about the nature of the Great Recession with the benchmark model should also be robust.

To examine this robustness directly, Figure 11 plots smoothed probabilities of two contractionary regimes over time for the key alternative cases considered in this section. In particular, we consider the general model reported in Table 4, the model for output per capita reported in Table 5, the model for different assumptions about structural change re-

Table 7: Estimates of parameters under alternative structural break assumptions

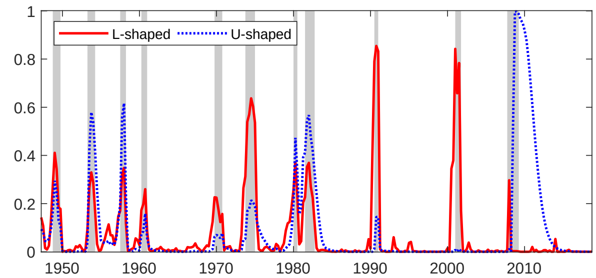
	No break		Break in 2000Q2		Dynamic Demeaning	
Parameter	Estimate	S.E.	Estimate	S.E.	Estimate	S.E.
p_{01}	0.015	0.011	0.020	0.012	0.031	0.015
p_{02}	0.020	0.011	0.017	0.010	0.019	0.011
p_{11}	0.650	0.206	0.666	0.153	0.744	0.102
p_{22}	0.734	0.122	0.724	0.129	0.734	0.120
μ_0	0.773	0.040	0.940	0.049	0.088	0.039
μ_1	-1.482	0.287	-1.545	0.213	-1.191	0.154
μ_2	-2.027	0.282	-2.100	0.291	-2.114	0.269
λ_2	0.405	0.056	0.420	0.058	0.423	0.054
δ			-0.388	0.071		
σ_{v0}	0.927	0.068	0.895	0.065	0.868	0.064
σ_{v1}	0.458	0.029	0.422	0.026	0.433	0.027
log-lik	-329.10		-315.05		-321.21	

Notes: The model is the same as the benchmark case in (3), except we assume no break in trend growth in the first case, the break in trend growth occurs in 2000Q2 in the second case, and model in (13) in the third case. For comparability to the benchmark model, we assume a structural break in residual volatility in 1984Q2 and set the length of the post-recession bounceback effect to $m = 5$. The standard errors are calculated using numerical second derivatives. We jointly estimate μ_2 and λ_2 using the restriction of $\mu_2 + m \cdot \lambda_2 = 0$, but report estimates and standard errors for both parameters..

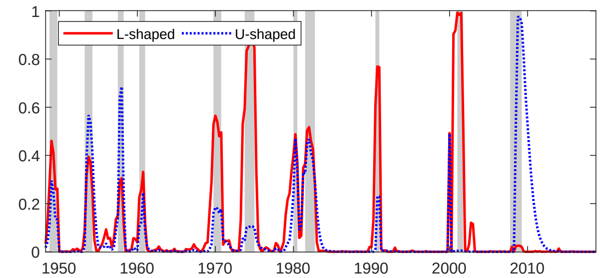
ported in Table 7, and the weighted-average approach discussed above. The classification of recessions changes across the different cases. For example, it is clear that considering the trend growth break in 2000 means that the 2001 recession would no longer be classified as a contractionary regime, supporting the idea that this timing for the structural break may be overfitting the temporary effects of the recession on growth rates. However, despite different inferences about some of the other recessions, the Great Recession is classified as being U shaped in all cases. Thus, we can be confident that our inferences about the nature of the Great Recession in particular are robust to different assumptions about structural change in trend growth.

Figure 12 directly confirms this robustness by reporting the estimated permanent level effects of different recessions for the various alternative cases. While the inferences about the 2001 recession in particular are quite sensitive to different assumptions, the consistent finding is that the Great Recession had comparatively little permanent effects on the level of output. Notably, this finding holds even under the assumption of no structural break in trend growth. Thus, even though it might be a surprising result given conventional views about

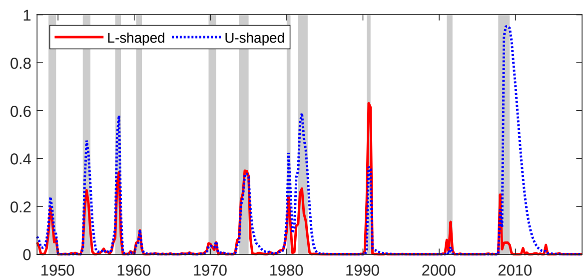
Figure 11: Smoothed probabilities of contractionary regimes for alternative cases



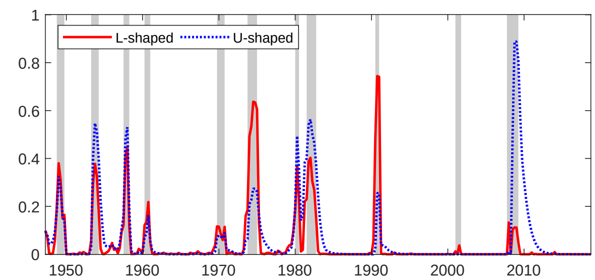
(a) General model



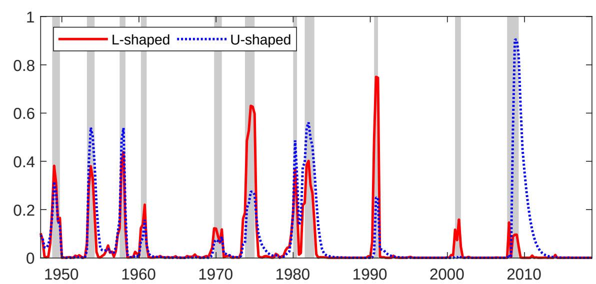
(b) Output growth per capita



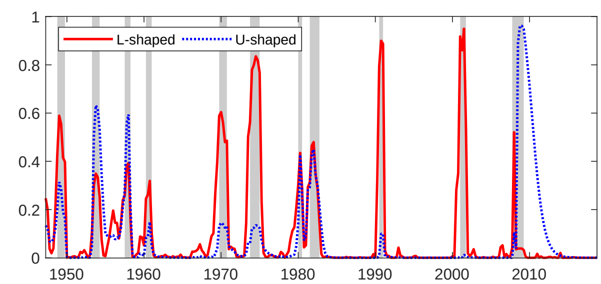
(c) No break



(d) Break in 2000Q2



(e) Weighted average over possible breakdates



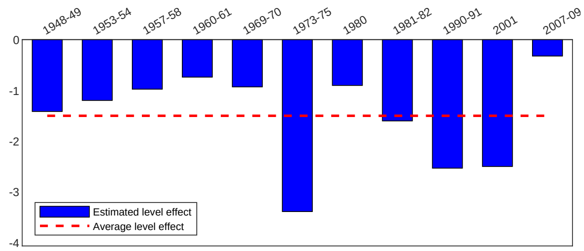
(f) Dynamic demeaning

Notes: The breakdate for residual volatility is given by $\tau_v = 1984Q2$. The shaded areas denote NBER recession dates.

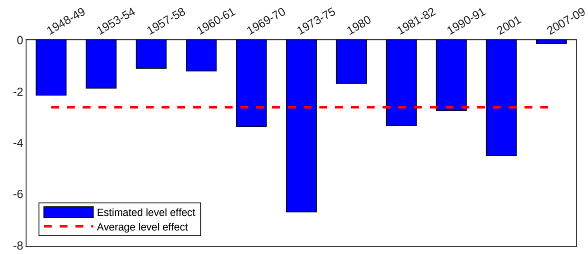
hysteresis effects of the Great Recession, our model makes the clear and robust inference that it did not, in fact, have large permanent effects on the level of output.

As with our benchmark model, the general implication is that the Great Recession corresponded to a large persistent negative output gap. This is illustrated in Figure 13, which plots log real GDP and the estimated trend around the Great Recession for the various alternative cases. In all cases, consistent with a persistent negative output gap, real GDP falls below trend in the Great Recession and remains well below trend for a number of years after the end of the recession.

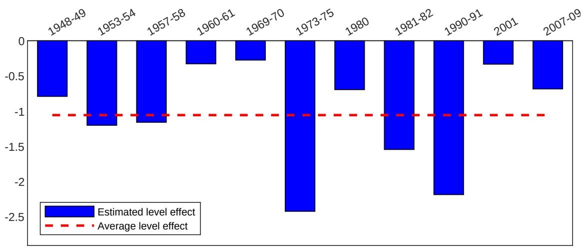
Figure 12: Level hysteresis effects of recessions for alternative cases



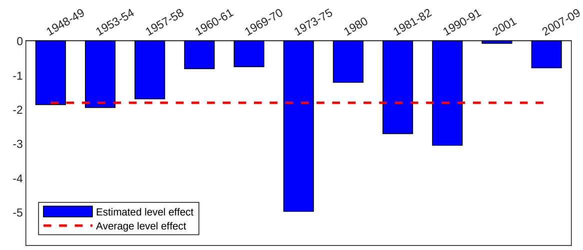
(a) General model



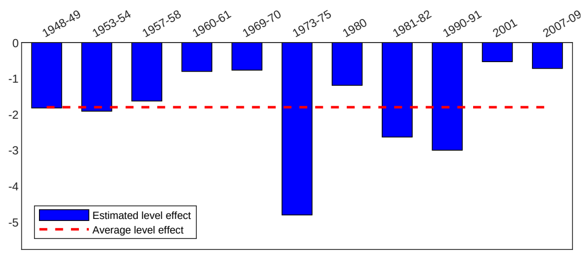
(b) Output growth per capita



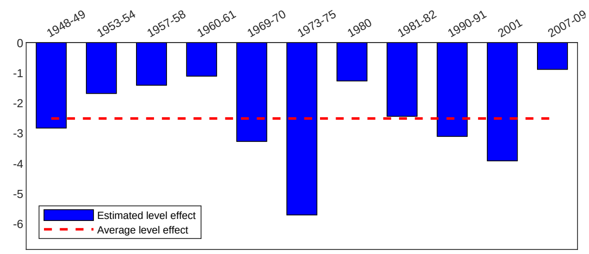
(c) No break



(d) Beak in 2000Q2



(e) Weighted average over possible breakdates



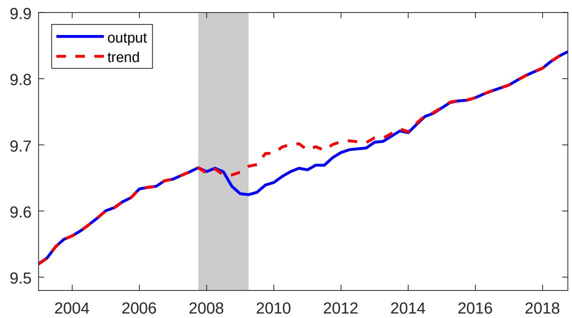
(f) Dynamic demeaning

Note: The breakdate for residual volatiltity is given by $\tau_v = 1984Q2$. The shaded areas denote NBER recession dates.

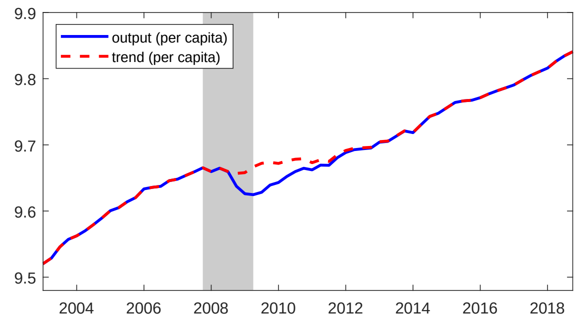
6 Conclusion

We have developed a new Markov-switching model of real GDP growth that accommodates two different types of recessions and allows for structural change in trend growth. Applying our model to U.S. data, we find that, perhaps surprisingly, the Great Recession was U shaped and did not appear to have hysteresis effects. Instead, the Great Recession generated a large persistent negative output gap, with the economy eventually recovering to a lower-growth trend path that, consistent with Fernald et al. (2017), appears to be due to a reduction in

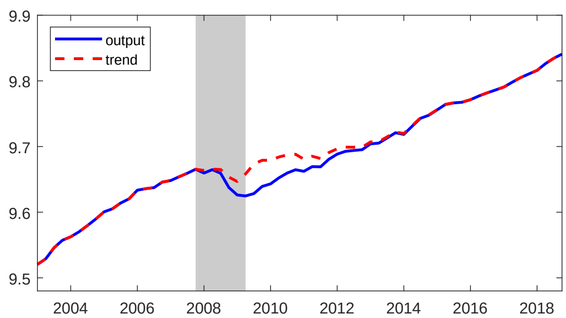
Figure 13: Output and trend around the Great Recession for alternative cases



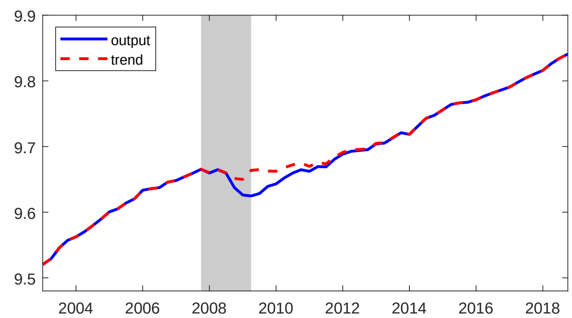
(a) General model



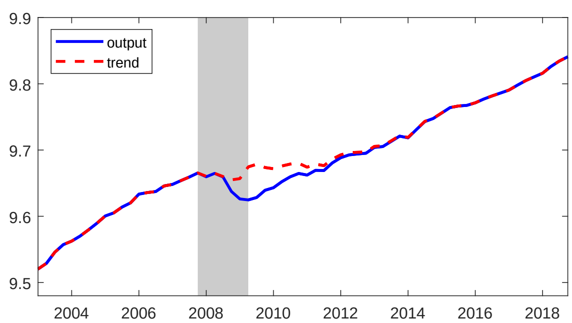
(b) Output growth per capita



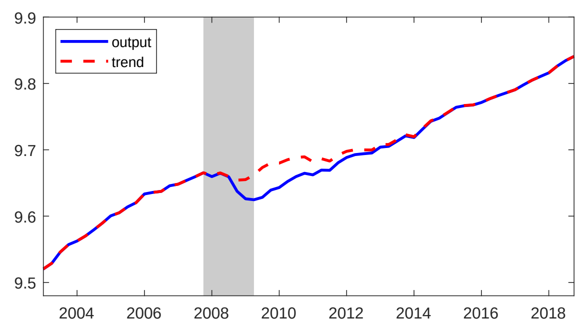
(c) No break



(d) Beak in 2000Q2



(e) Weighted average over possible breakdates



(f) Dynamic demeaning

Note: The shaded areas denote NBER recession dates.

productivity growth that began no later than 2006. We highlight that our inferences about the timing of the output growth slowdown are sharpened by our consideration of a time series model that accounts for nonlinear dynamics of recessions. Meanwhile, our inferences about the nature of the Great Recession as generating a persistent negative output gap rather than large hysteresis effects is highly robust to different assumptions regarding the nature of structural change in trend growth.

Our analysis is univariate and we leave consideration of the implications of our findings for a multivariate setting to future research. However, we note that, similar to the conclusions in Huang and Luo (2018), our estimated output gap can clearly help explain weak inflation in the years immediately after the Great Recession. Our results also suggest that the slow growth of the U.S. economy is likely to persist, despite interest rates no longer being at the zero-lower-bound.

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