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Did Okun's Law Die after the Great Recession?

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Abstract: This paper proposes an empirical framework to estimate Okun's law which focuses on structural breaks and threshold nonlinearity. We use sequentially the Bai and Perron’s (1998, 2003) structural break and threshold methodology to enable regime-dependent as well as threshold-dependent changes in the unemployment rate. We employ an autoregressive distributed lag version of Okun's law in first differences, which allows for delayed reactions of the unemployment rate to output changes. Applied to US data (1948Q1-2015Q4), the empirical analysis characterize Okun’s law as a three-regime relationship with the first break coinciding with the 1973 oil shock, and the second break immediately following the end of the Great Recession. In the post-Great Recession regime, we find that Okun’s law breaks down as a linear relationship. This result assumes a linear and symmetric relationship between changes in the unemployment rate and real output. We test this assumption for each of the identified regimes using threshold estimation and recognize a threshold within each regime, which rejects the linearity and symmetry hypotheses and, thus, suggests that Okun’s law follows a more complex nonlinear asymmetric dynamics. Importantly, when we apply threshold estimation to the post-Great Recession regime, we find that the time-honored link between output growth and the unemployment rate still holds.

Keywords: Okun's law; structural breaks; threshold effects.

JEL classification: C14; E31; C22

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1. Introduction

Since Okun's (1962) seminal contribution, a large literature documents substantial evidence of the correlation between changes in real output and changes in the unemployment rate, an enduring relationship that has become known in the macroeconomic literature as Okun's law.[1]

An extensive body of research has evaluated the strength of Okun's law across countries and time periods. The relationship has been estimated either adopting a single country approach (Clark, 1983; Gordon, 1984; Adams and Coe, 1989; Prachowny, 1993; Blinder, 1997; Attfield and Silverstone, 1998), a panel-data approach (Fouquau, 2008; Huang and Yeh, 2013; Freeman, 2001) or a multi-regional approach (Huang and Yeh, 2013; Adanu, 2005; Apergis and Rezitis, 2003; Villaverde and Maza, 2009). In the conventional macroeconomic landscape, Okun's law has been accepted as an empirical regularity. Theoretically, the relationship is viewed as the link between the aggregate demand and the Phillips curve; empirically, the "Okun's coefficient" is considered a useful "rule of thumb" in economic forecasting and policy modeling.

Recent research, however, has questioned the robustness of the relationship. Lee (2000) documents substantial disparities between the estimates for the US and many European countries, arguably a consequence of hysteretic effects in European labor markets. Besides cross-country heterogeneity, the robustness of the estimates has been explored by focusing on the stability of the parameters and the linearity of Okun's law. On the one hand, a growing literature dealing with Okun's law questions the structural stability of Okun's law. Knotek (2007), Apergis and Rezitis (2003), Lee (2000), Moosa (1997), Sogner and Stiassny (2002), Huang and Chang (2005), Huang and Lin (2006), Schnabel (2002), Owyang and Sekhposyan (2012), and

[1] This empirical relationship forms a major part of every traditional macro-model as the aggregate supply curve comes from combining Okun's law with the Phillips curve. Moreover, this relationship also leads to important implications for macroeconomic policy. First, it documents what rate of growth of output leads to a reduction in the unemployment rate. Second, the effectiveness of disinflation policy depends on the responsiveness of unemployment on the rate of output growth.

Osterholm (2016) find strong evidence of structural change in Okun's law.[2] On the other hand, another growing literature dealing with Okun's law questions the linear nature of the relationship. Lee (2000), Viren (2001), Harris and Silverstone (2001), Holmes and Silverstone (2006), Mayers and Viren (2002), Crespo-Cuaresma (2003), Huang and Chang (2005), Knotek (2007), Silvapulle et al. (2004), Owyang and Sekhposyan (2012), and Chinn, Ferrara and Mignon (2014) find evidence that Okun's law conforms to a nonlinear and asymmetric relationship.

Macroeconomic forecasting could benefit from a better understanding of structural breaks and nonlinearities in Okun's law, resulting in a decrease in forecasting errors. Stabilization policies designed to mediate the effect of output on the unemployment rate during recessions could benefit from an understanding of the threshold nonlinearities in Okun's law. Okun's law provides policy-makers with a benchmark to measure the relative cost of output in terms of the unemployment rate. An incorrectly estimated benchmark could lead to policy mistakes.

The major limitation of the literature dealing with structural breaks and nonlinearities in Okun's law, with the exception of Owyang and Sekhposyan (2012), Chinn, Ferrara and Mignon (2014), and Osterholm (2016), is that the sample period does not extend beyond 2002 and, hence, does not reflect the period when Okun's law has been arguably under the most intense scrutiny, which includes the onset of the Great Recession, its aftermath, and the recovery that began in the mid-2009.

Whether the Great Recession (2007-2009) led to fundamental structural changes in the US economy and, consequently, in the validity of Okun's law, is critically important. Within this

[2] In contrast, Ball, Leigh and Lougani (2015) estimate Okun's law for a sample of 20 advanced countries using annual data from 1989 to 2012 and show that there is a strong and stable relationship between output and unemployment. Similarly, Sogner (2001), using a Markov switching methodology, finds that in the Austrian economy Okun's law is a stable relationship.

context, one of the puzzles of the post-Great Recession period has been the behavior of the unemployment rate and real output growth. For example, real GDP growth contracted by half a percentage point during 2009, yet the unemployment rate jumped by 3.0 percentage points. If Okun's law had held in 2009, the unemployment rate would have risen by about half as much as it did over the course of the year (Daly and Hobijn, 2010). If the relationship in 2009 was unusual in one direction, in 2011 it was unusual in the opposite direction, with the unemployment rate falling from 9.1 percent to 8.3 in 2011, while real GDP grew only 1.6 percent. The puzzle of the "jobless" recovery of 1991 and 2001 (Holmes and Silverstone, 2006; Gordon, 1993), an economic recovery that fails to create jobs, gave way to the puzzle of the "GDP-less" recovery of 2011 (Gordon, 2015), a recovery that creates jobs in an anemic growth environment. The rate of unemployment rose more than Okun's Law predicted during the Great Recession and has fallen more than the law predicted since the Great Recession.

Did the connection between labor market and output market break down? There is no doubt that Okun's law has been an important paradigm in macroeconomics and a powerful "rule of thumb" in economic modeling and forecasting. But the validity of any "rule of thumb" impinges upon the stability, linearity, and symmetry of its construct. An unstable, non-monotonic "rule of thumb," that does not follow a single pattern, is not much of a rule (Mayer and Tasci, 2012).

In this paper, we analyze the dynamics of Okun's law using the most updated dataset as of now (1948Q1-2015Q4) and give some insight into the shifting patterns of Okun's law over time and its modified nature during the "GDP-less" recovery of post-2009. We propose an empirical framework to estimate Okun's law that takes into account structural breaks and threshold nonlinearity simultaneously. Previous research generally treats the temporal instability of Okun’s law and the threshold nonlinearity of the relationship as separate issues. Ideally, however, these two issues should be examined in a unified framework. In this paper, we model Okun’s law using sequentially the Bai and Perron’s (1998, 2003) structural breaks and threshold methodology, which enable regime-dependent as well as threshold-dependent changes in the unemployment rate. Applied to US data over the period 1948Q1-2015Q4, our empirical findings characterize Okun’s law as a three-regime nonlinear asymmetric relationship, where each regime is characterized by a differential threshold. We identify the first date break with the 1973 oil shock, and the second date break with the aftermath of the Great Recession. Without consideration of the threshold structure embedded in the relationship, we find that Okun’s law breaks down in the third regime. Importantly, however, when the threshold analysis is conducted within each regime, we identify a threshold in the third regime, where the time-honored link between output growth and change in the unemployment rate still holds, albeit in a nonlinear and asymmetric form.

This analysis has not been carried out previously for Okun's law in the US. This new scope of analysis offers interesting results that analysts and policy makers should incorporate into their policy analysis. One key problem that arises when dealing with the Okun’s law is that several specifications co-exist in the empirical literature, rendering comparisons difficult. Specifically, Okun’s law has received two main specifications -- the gap and difference versions. The gap version identifies the empirical regularity between the output gap -- the difference between actual real GDP and capacity output -- and the unemployment rate gap -- the difference between the actual unemployment rate and the natural rate of unemployment. The difference version considers the empirical regularity between the growth rate of output and the change in the unemployment rate. The gap version requires the determination of capacity output and the natural unemployment rate, which are not observable. The difference version, conversely, applies to observed data on output and rate of unemployment. In this paper, we adopt the difference version of Okun’s law due to its simplicity, accuracy, and direct applicability to the original data and not require strong, and sometimes controversial, assumptions regarding the definition and computation of potential output and full employment. Among others, the difference version represents a convenient way to achieve stationarity when the unemployment rate and real GDP contain a unit root.

The remainder of the paper is organized as follows. Section 2 outlines the empirical methodology, namely the Bai-Perron (1998, 2003) structural break tests and threshold estimation of the Bai-Perron regimes. Section 3 displays the results of the empirical analysis. We present three sets of empirical results. First, we present the results from the constant parameter linear model, assuming no breaks and no thresholds within the breaks. Second, we present the results from the structural break model without threshold. Finally, we present the results from the structural break model with threshold. In other words, we first investigate whether the model exhibits structural breaks. Finding two structural breaks and, thus, three regimes, we then examine each regime for evidence of threshold nonlinearity. Section 4 concludes.

2. The empirical models

Okun's law, in its difference version, postulates that the growth rate in real gross domestic product (GDP) drives the change in the unemployment rate. This assumes that an increase in output will require more factor input leading to a lower unemployment rate. The standard difference version, written as a linear regression model, is given by:

$$\Delta U_t = \alpha_0 + \alpha_1 \Delta GDP_t + \varepsilon_t , \tag{1}$$

where $\Delta U_t = U_t - U_{t-1}$ and U_t is the rate of unemployment, measured in percentage,

$\Delta GDP_t = GDP_t - GDP_{t-1}$ and GDP_t is real gross domestic product, measured in natural logarithms, and ε_t is the error term. The coefficient α_0, a constant term, measures the change in the unemployment rate when growth is null. The parameter α_1 measures the effect on the unemployment rate of a 1 percent change in the growth rate in real GDP. This parameter, or "Okun's coefficient," has a negative value, capturing the negative relationship between the unemployment rate and GDP growth. Thus, if the economy increases its growth rate by 1 percent, the unemployment rate will change by α_1 percent. Using quarterly data from 1947Q2 to 1960Q4, on the unemployment rate and real gross national product (GNP), Okun (1962) estimated α_0= 0.3 and α_1= -0.3, leading to the conclusion that a 1-percent increase in real GNP growth associated with a 0.3 percentage point decrease in the unemployment rate. The ratio $-(\alpha_0 / \alpha_1)$ gives the growth rate at which the unemployment rate is stable, except for random shocks.

This "static version" of Okun's law (Knotek, 2007), captures only the contemporaneous correlation, and ignores the rich dynamics between ΔU_t and ΔGDP_t, such as the effect of past GDP growth on the current unemployment rate or the effect of the past unemployment rate on the current unemployment rate, as suggested by the persistence of the unemployment rate literature (Barro, 1988; Mortensen and Pissarides, 1994) and the hysteresis hypothesis (Blanchard and Summers, 1987; León-Ledesma, 2002) . Knotek (2007) shows that the business cycle and variation in the timing of the connection between growth and the unemployment rate affects the version of Okun's law in equation (1). One argument favors including past GDP growth to capture the idea of "jobless recoveries". If true, then after a recession, the recovery of employment lags the recovery of output and, thus, both employment and the unemployment rate will not only depend on current output but also on past output values. Including past changes in the unemployment rate on the right-hand side of the dynamic version of Okun’s law helps to eliminate serial correlation in the error terms that comes from regressing the differences (Weber, 1995; Moosa, 1997). For these reasons, we specify Okun's law within a simple autoregressive distributed lag framework (see Weber, 1995; Sogner and Stiassny, 2002) as follows:

$$\Delta U_t = \alpha_0 + \alpha_1 \Delta GDP_t + \alpha_2 \Delta GDP_{t-1} + \alpha_3 GDP_{t-2} + \alpha_4 \Delta U_{t-l} + \varepsilon_t . \qquad (2)$$

In this "dynamic version" of Okun's law (Knotek, 2007; Owyang and Sekhposyan, 2012),[3] the coefficient α_1 measures the contemporaneous effect of output growth, whereas the sum $\alpha_1 + \alpha_2 + \alpha_3$ measures the total short-run effect. This specification allows us to calculate the long-run effect of output growth on the unemployment rate change as follows:

$$\theta = \frac{\alpha_1 + \alpha_2 + \alpha_3}{1 - \alpha_4} . \qquad (3)$$

In addition to the short- and long-run dynamics of the link between the unemployment rate and output growth, the empirical literature has been increasingly concerned with the presence of discrete changes (structural breaks) and threshold nonlinearity and asymmetry of the relationship.

Okun's law in either the static or dynamic versions assumes linearity and symmetry: expansions and contractions in output exert the same absolute effect on the unemployment rate. As mentioned by Lundbergh, Terasvirta and van Dijk (2003), ample empirical evidence exists for both structural breaks and nonlinearity in the dynamic properties of many macroeconomic series. Gordon (1984) and Evans (1989) suspect that the unemployment rate and output dynamics underwent structural change following the crude oil shocks in the 1970s, resulting in a

[3] The distributed lag specification reduces the simultaneous equation bias for the total effect, as long as output growth is positively autocorrelated (Sogner and Stiassny, 2002).

structural break in Okun's relationship. Lee (2000) attributes the developments to structural changes caused by corporate restructuring, productivity and wages slowdown, and rising female labor-force participation. Moreover, Harris and Siverston (2001) argue the importance of testing for asymmetry and nonlinearity in the relation between the unemployment rate and output. Among other reasons, if Okun's law is asymmetric, then the Phillips curve is also probably asymmetric. Moreover, how the unemployment rate reacts to changes in output has implications for the labor market and the appropriate monetary and fiscal policies responses. Viren (2001) argues that asymmetry and nonlinearity in Okun's law can result in varying degrees of effectiveness of unemployment policies. Conceptually, the existence of asymmetries in the relationship between the unemployment rate and output growth does not necessarily invalidate Okun’s law. Although the original specifications impose symmetric responses, symmetry may prove non-optimal, as no theoretical reason exists to justify the presence of symmetric responses in the relationship between the unemployment rate and output growth.

To test the model for structural breaks and threshold effects, we proceed sequentially in two basic steps. First, following the Bai and Perron (1998, 2003) methodology, we test the “dynamic” version for multiple structural breaks or regime shifts. Within this framework, assuming *m* breaks (*m*+1 regimes), equation (2) becomes

$$\begin{aligned}
&\Delta U_t = \alpha_0^1 + \alpha_1^1 \Delta GDP_t + \alpha_2^1 GDP_{t-1} + \alpha_3^1 \Delta GDP_{t-2} + \alpha_4^1 \Delta U_{t-1} + \varepsilon_t \\
&\text{for } t = 1, \ldots, T_1 \\
&\Delta U_t = \alpha_0^2 + \alpha_1^2 \Delta GDP_t + \alpha_2^2 \Delta GDP_{t-1} + \alpha_3^2 \Delta GDP_{t-2} + \alpha_4^2 \Delta U_{t-1} + \varepsilon_t \\
&\text{for } t = T_1 + 1, \ldots, T_2 \\
&\ldots \\
&\Delta U_t = \alpha_0^m + \alpha_1^m \Delta GDP_t + \alpha_2^m \Delta GDP_{t-1} + \alpha_3^m \Delta GDP_{t-2} + \alpha_4^m \Delta U_{t-1} + \varepsilon_t \\
&\text{for } t = T_m + 1, \ldots, T
\end{aligned} \tag{4}$$

where the superscript refers to the break. The indices T_1, ... ,T_m are the break points, which are explicitly treated as endogenous and unknown. The system of equations in (4) indicates, in the terminology of Bai and Perron (1998, 2003), a pure structural-change model because all coefficients can change. A partial structural-change model, on the other hand, arises if some of the coefficients cannot change, that is, are estimated over $t = 1,...,T_1$. The Bai-Perron procedure consists of estimating the unknown regression coefficients together with the break points. Bai and Perron (1998, 2003) suggest obtaining the estimators via the least squares approach and provide a dynamic programming algorithm to efficiently compute the estimates. In addition, they also propose three formal *F*-related test statistics for multiple breaks, namely the $\sup F_T(\ell)$ test, the double maximum tests (UD max and WD max), and sequential tests $\sup F_T(\ell+1|\ell)$. The $\sup F_T(\ell)$ statistic tests the null hypothesis of no structural break (ℓ = 0) against the alternative of a fixed (arbitrary) number of breaks (ℓ = 1,…, *k*). The two maximum statistics test the null hypothesis of no structural break against the alternative of an unknown number of breaks, given some upper bound *M*, $1 \le \ell \le M$. The UD and WD max are equal-weight and weighted versions, respectively, where the weights depend on the number of regressors and the significance level of the test. In both tests, the break points are estimated using the global minimization of the sum of squared residuals. The $\sup F_T(\ell+1|\ell)$ statistics, on the other hand, sequentially test the null hypothesis of ℓ breaks against the alternative of ℓ + 1 breaks, ℓ = 1,2, … . According to simulation results, a useful strategy for selecting the number of breaks first considers the UD max or the WD max tests to see if at least one break exists. If these tests indicate the presence of at least one break, then the number of breaks can be decided based on a sequential determination of the $\sup F_T(\ell+1|\ell)$ statistics, which involves choosing the number of breaks *m* such that

$\sup F_T(\ell+1|\ell)$ are insignificant for $\ell > m$. Bai and Perron (2003) argue that this approach leads to the best results and they recommend it for empirical applications. Please refer to Bai and Perron (1998, 2003) for technical details. Hansen (2001) and Perron (2006) offer useful overviews of the structural breaks literature.

Second, within each regime identified by the Bai-Perron procedure, we test the "dynamic" version for threshold nonlinearity. That is, we apply threshold estimation, which can capture complex nonlinearities and complex dynamics described by observed variables crossing unknown thresholds. Letting q_{t-d} be the threshold variable, where d is the delay parameter, which is determined endogenously together with its associated threshold parameter ξ, and assuming, for simplicity, a single threshold in a model with m structural breaks or m+1 regimes, we can express the linear equation (2) as a non-linear equation under a two-state threshold model as follows:

$$
\begin{aligned}
&\Delta U_t = \alpha_{0,1}^1 + \alpha_{1,1}^1 \Delta GDP_t + \alpha_{2,1}^1 GDP_{t-1} + \alpha_{3,1}^1 GDP_{t-2} + \alpha_{4,1}^1 \Delta U_{t-1} + \varepsilon_t \\
&\text{if } q_{t-d} < \xi^1 \text{and } t = 1,\dots,T_1 \\
&\Delta U_t = \alpha_{0,2}^1 + \alpha_{1,2}^1 \Delta GDP_t + \alpha_{2,2}^1 GDP_{t-1} + \alpha_{3,2}^1 \Delta GDP_{t-2} + \alpha_{4,2}^1 \Delta U_{t-1} + \varepsilon_t \\
&\text{if } q_{t-d} \geq \xi^1 \text{and } t = 1,\dots,T_1 \\
&\Delta U_t = \alpha_{0,1}^2 + \alpha_{1,1}^2 \Delta GDP_t + \alpha_{2,1}^2 \Delta GDP_{t-1} + \alpha_{3,1}^2 \Delta GDP_{t-2} + \alpha_{4,1}^2 \Delta U_{t-1} + \varepsilon_t \\
&\text{if } q_{t-d} < \xi^2 \text{ and } t = T_1 + 1,\dots,T_2 \qquad (5) \\
&\Delta U_t = \alpha_{0,2}^2 + \alpha_{1,2}^2 \Delta GDP_t + \alpha_{2,2}^2 \Delta GDP_{t-1} + \alpha_{3,2}^2 \Delta GDP_{t-2} + \alpha_{4,2}^2 \Delta U_{t-1} + \varepsilon_t \\
&\text{if } q_{t-d} < \xi^2 \text{ and } t = T_1 + 1,\dots,T_2 \\
&\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots\dots \\
&\Delta U_t = \alpha_{0,1}^m + \alpha_{1,1}^m \Delta GDP_t + \alpha_{2,1}^m \Delta GDP_{t-1} + \alpha_{3,1}^m \Delta GDP_{t-2} + \alpha_{4,1}^m \Delta U_{t-1} + \varepsilon_t \\
&\text{if } q_{t-d} < \xi^m \text{and } t = T_m + 1,\dots,T
\end{aligned}
$$

$$\Delta U_t = \alpha_{0,2}^m + \alpha_{1,2}^m \Delta GDP_t + \alpha_{2,2}^m \Delta GDP_{t-1} + \alpha_{3,2}^m \Delta GDP_{t-2} + \alpha_{4,2}^m \Delta U_{t-1} + \varepsilon_t$$

$$\text{if } q_{t-d} < \xi^m \text{ and } t = T_m + 1, ..., T$$

where the second subscript refers to the threshold. In this formulation, the coefficients of the "dynamic" version of Okun's law depend upon the regime, and within a given regime, upon the threshold. The threshold regression model implies that the regression parameters differ depending on the value of threshold variable. When the threshold variable falls below the threshold parameter, the model estimates the first equation of the regime, while when the threshold variable exceeds the threshold parameter, the model estimates the second equation of the regime. We apply the Bai-Perron procedure to determine the number of thresholds, rather than the fixed regressor bootstrapped testing proposed by Hansen (1999). Bai and Perron (2001) show that estimation of the threshold and breakpoint models are fundamentally equivalent.[4] Accordingly, we can generally apply the three formal *F*-related test statistics for multiple breaks, namely the $\sup F_T$ test, the double maximum tests (UD max and WD max), and sequential tests $\sup F_T(\ell + 1|\ell)$ in the context of threshold regression. We can, thus, employ the global $\sup F_T(\ell)$ test and the double maximum tests (UD max and WD max) to detect whether at least one threshold exists, while we can use the sequential $\sup F_T(\ell + 1|\ell)$ tests to detect the number of thresholds. If the threshold variable falls below the threshold parameter, then we estimate equation (4), while if it exceeds the threshold parameter, then we estimate equation (5). The threshold regression model implies that the regression parameters differ depending on the value of threshold variable

3. Data and empirical results

[4] Threshold regressions can be thought of as breakpoint OLS regressions with data reordered with respect to the threshold variable. Alternatively, breakpoint regressions may be thought of as threshold regressions with time as the threshold variable.

We employ quarterly data on the real GDP growth rate (not annualized) and the quarterly average of the monthly civilian unemployment rate from 1948Q1 through 2015Q4. The data come from FRED at the Federal Reserve Bank of St. Louis (https://research.stlouisfed.org/fred2/). Prior to presenting the results, we determine whether ΔU_t and ΔGDP_t contain unit roots. First, we implement a linear ADF and PP tests with a constant term to determine the order of integration of ΔU_t and ΔGDP_t. We select the optimum lag length via the Schwarz information criterion (SIC). The SIC indicates in both cases that 0 lags render white-noise residuals in the ADF and PP tests. The relevant test statistics easily reject the null hypothesis of a unit root in both ΔU_t and ΔGDP_t, using both the ADF and PP tests. This ensures stationarity of both series, a necessary condition for valid Bai-Perron tests.

Before performing the econometric analysis, we first consider simple scatter charts for the growth rate of real GDP and the change in the unemployment rate and trend lines. This provides a preliminary data analysis to set the stage for our econometric analysis that follows. Figure 1 plots the scatter chart using quarterly data from 1949Q1 through 1960Q4, which corresponds to Okun's original sample period. A 1-percent decrease in the real GDP growth rate translates into a 0.3321 percent increase in the unemployment rate. Figure 2 extends the descriptive analysis to cover the entire sample period from 1949Q1 through 2015Q4. Now, a 1-percent decrease in the real, GDP growth rate leads to a 0.2825 percent increase in the unemployment rate. Finally, the Great Recession caused the biggest disruption to the macroeconomy in the post WWII period. Thus, Figure 3 plots the scatter chart where we use subsamples of 1949Q1 to 2006Q4, 2007Q1 to 2009Q4, and 2010Q1 to 2015Q4. The response coefficient for the pre-Great Recession and Great Recession samples are -0.2849 and -0.3397, respectively. The response coefficient for post-Great Recession is -0.0155, or nearly no response to changes in the unemployment rate. Note, also, that the response coefficient increased during the Great Recession relative to the pre-Great Recessions response coefficient.

3.1. Results from the linear constant parameter model

For ease of exposition, we first present the empirical results for the linear constant-parameter autoregressive distributed lag as in equation (2), which is the model to which the structural break model collapses if the regression parameters are equal across the regimes.

Table 1 reports the OLS findings with standard errors corrected for heteroskedasticity and autocorrelation using the Newey-West (1987) covariance matrix. Table 1 shows that the intercept is positive and significantly different from zero at the 1 percent level. The contemporaneous effect, Okun's coefficient on ΔGDP_t, is -0.204 and significantly different from zero at the 1 percent level. The total short-run effect, the sum of the estimates $\hat{\alpha}_1 + \hat{\alpha}_2 + \hat{\alpha}_3$ is -0.293 exceeds the contemporaneous effect in absolute value and significantly differs from zero at the 1 percent level. This value does not differ much from the findings of Okun (1962), Moosa (1997), and Crespo-Cuaresma (2003). The estimate of α_4 of 0.315 is significant at the 1 percent level and positive, suggesting a positive dependence on the previous quarter unemployment rate and a monotonic adjustment toward equilibrium. Since the estimate of α_4 is less than unity, the long-run estimates of Okun’s relationship are significant at the 1-percednt level -- $\frac{\hat{\alpha}_0}{1-\hat{\alpha}_4}$ = 0.332 and $\frac{\hat{\alpha}_1 + \hat{\alpha}_2 + \hat{\alpha}_3}{1-\hat{\alpha}_4}$ = -0.428. In the long-run, a stronger relationship exists between the unemployment rate and output growth, as a 1-percent increase (decrease) in real GDP growth associates with a -0.428 percent decrease (increase) in the rate of unemployment. Thus, overall, the OLS results of the dynamic version of Okun’s law provide strong evidence in support of the inverse linear relationship between the unemployment rate and GDP growth.[5] Some evidence of residual serial correlation at longer lags exists. The Ramsey Reset Test indicates possible nonlinearities, since we can reject the null hypothesis of linear specification at 5-percent level.

3.2. Results from the structural break model without threshold

Do structural breaks exist in the relationship between changes in the unemployment rate and output growth? Do the estimates of the Okun coefficient change in the presence of structural breaks? Testing for structural change has become an important issue in econometrics because a multitude of political and economic factors can cause the relationships among economic variables to change over time. This is particularly relevant for the relationship between the unemployment rate and output growth. Early work on structural breaks (e.g., Quandt, 1958; and Chow, 1960) focused on testing for structural change at a single known break date. More recently, however, the econometric literature has developed methods that allow estimating and testing for structural change at unknown break dates (e.g., Andrews, 1993; and Andrews and Ploberger, 1994) for the case of a single structural break, and Bai and Perron (1998, 2003a, 2003b) for the case of multiple structural changes. In this section, we estimate Okun's law through a multiple endogenous break model, making use of the approach of Bai and Perron (1998, 2003). The Bai-Perron procedure allows testing for multiple breaks at unknown dates so that we estimate each successive break point using a specific-to-general strategy in order to determine consistently the number of breaks. Table 2 contains the Bai-Perron tests of structural breaks applied to the dynamic version of Okun's law.

[5] For comparison, we also estimate the static version of Okun's law. The significant estimates of the intercept and Okun's coefficient are 0.226 and -0.285, respectively. These estimates nearly match the long-run estimates obtained with the dynamic version. The adjusted R-squared, however, is only 0.469, which is 30 percent lower than the adjusted R-squared for the dynamic version. Moreover, the Ljung-Box statistics at lags 4, 8, and 10 are, respectively, 36.773, 38.141, and 38.575 ,which all indicate significant residual serial correlation.

We trim by 10 percent, so that each subsample contains at least approximately 26 observations (i.e., about seven years of data), and allows for a maximum of 5 breaks. We also incorporate heterogeneous error distributions across breaks. At the same time, since we include a lagged dependent variable as a regressor, we rule out serial correlation in the errors (see Assumption A4 in Bai and Perron, 1998). The $\sup F_T(\ell)$ tests uniformly reject at the 5 percent level the null hypothesis of no structural break against the alternative of ℓ breaks (ℓ = 1…5). The double maximum test statistics, UD max and WD max, confirm this finding. The UD max and WD max statistics are significant at the 5 percent level, which indicates the presence of at least one break. Thus, both the $\sup F_T(\ell)$ and the UD max and WD max statistics provide strong evidence of at least one structural break. The sequential test statistics $\sup F_T(\ell+1|\ell)$, on the other hand, are not significant for $\ell > 2$ suggesting a modelwith two breaks only.

Combined with the previous results, the overwhelming and consistent evidence suggests a model with three regimes, with the endogenous break dates estimated at 1973Q2 and 2009Q3. The first break date coincides with the first oil shock, while the second date break coincides with the start of the post-Great Recession period. Our findings differ from Weber (1995), who finds no indication of structural break in 1973. Importantly, we do not find that the Great Recession represents a break date of Okun's law. Interestingly, the 2009Q3 break date coincides with the one break date estimated using the static version of Okun's law. Moreover, it also coincides with the one date break found in ΔU_t, implying that the Okun's regressions from equations (1) and (2) reflect the same structural change as the dependent variable.[6]

To check further the robustness of our results, we also apply the Quandt-Andrew test to the residuals of the static and dynamic versions of Okun's law. This one-break test does not

[6] We do not report the results, but will make them available on request.

assume a given break date. In the dynamic version, the maximum LR *F*-statistic equals 6.837, which is significant at hte1-percent level.. In the static version, the maximum LR *F*-statistic equals 8.795, which is also significant at the 1-percent level. In both cases, the break date is 2009Q2.

Given the strong evidence of structural breaks in 1973Q2 and 2009Q3, we now explore the characteristics of the three regimes. We separate the whole period into three regimes where the first regime is 1949Q2-1973Q2 (with 96 observations), the second regime is 1973Q3-2009Q2 (with 144 observations), and the third regime is 2009Q3-2015Q4 (with 26 observations), respectively, and reestimate the dynamic version of Okun's laws in each regime separately. Table 3 reports the OLS estimates of the three regimes. The estimates of the first regime, 1949Q2-1973Q2, are highly significant and possess the correct sign. Overall, these estimates do not differ much from the corresponding OLS constant parameter estimates in Table 1. The contemporaneous effect approximately equals the contemporaneous effect obtained under the assumption of no breaks. But the total effect, computed as a sum of $\alpha_1 + \alpha_2 + \alpha_3$, equals -0.384, which exceeds the corresponding value obtained under the assumption of no breaks. The long-run effect of ΔGDP_t on ΔU_t is -0.443, which nearly matches numerically the corresponding constant parameter estimates.

The results for the second regime, 2009Q3-2015Q4, maintain approximately the same pattern as the estimates of the first regime. The contemporaneous effect of output growth on the unemployment rate in the second regime nearly matches the contemporaneous effect in the first regime and the corresponding constant parameter estimate. The total effect in the second regime is approximately the same as that in the first regime, but the persistence of the unemployment rate, a measure of the speed of adjustment to equilibrium, exceeds that of the first regime.

Consequently, in the second regime, the long run effect of ΔGDP_t on ΔU_t is -0.487, which slightly exceeds that of the first regime.

The results for the third regime, 2009Q3-2015Q4, however, show strikingly that the explanatory variables play no significant role in Okun's relationship. The contemporaneous and lagged coefficients on ΔGDP_t do not significantly differ from zero and their sum , $\alpha_1 + \alpha_2 + \alpha_3$, which defines the total effect of output growth of -0.221,which proved insignificant. Furthermore, the long-run effect of ΔGDP_t on ΔU_t is -0.243, which does not significantly differ from zero at the 5-percent level. These findings underscore the irrelevance of inference when the estimated model does not consider the possibility of structural breaks.

It, thus, appears that in the post-Great Recession period, Okun's law does not hold either in the short-run or in the long-run. Although the fit of Okun's law in the first and second regimes remains good, the relationship appears to breaks down as a linear relationship immediately after the Great Recession, and the reason seems concentrated on 2009Q3. The Ljung-Box $Q(m)$ statistics at lags 4, 8, and 12 are 6.121, 9.876, and 18.121, respectively, indicating no residual autocorrelation at the 10-percent level.

To check the robustness of our results, we consider other methods of identifying structural breaks. The tests for breaks in all recursively determined partitions (Bai, 1997) and the tests based on global optimization procedures (Bai and Perron, 1998) confirm the number of breaks and the break dates identified by the sequential procedure. These results, therefore, strengthen the conclusions of the $\sup F_T(\ell+1|\ell)$ sequential tests. The only procedures that make a difference use the information criteria suggested by Yao (1998) and Liu, Wu, and Zidek (1997). The information criteria provide an alternative approach to inference on structural breaks for linear models. The implementation of information criteria approaches requires the search for the global minimum of the residual sum of squares, for which Bai and Perron (2003) developed an efficient search algorithm. Yao (1998) proposes using the Bayesian Information criterion (BIC), while Liu, Wu and Zidek (1997) recommend a modified Schwarz criterion (LWZ). Applying these criteria to equation (2), we find that the BIC selects only one break in 2009Q3, while the LWZ selects no break. Bai and Perron (2006), based on extensive simulations results, conclude that the BIC performs well unless the error term of the linear regression suffers from serial correlation, while the LWZ works well if the true model does not contain a structural break, but does not do well if the true model contains a structural break. Furthermore, because these information criteria cannot incorporate heterogeneity across regimes, Bai and Perron (2006) argue that the sequential procedure works best in determining the number of breaks.

3.3. Results from the structural break model with threshold

Does the presence of threshold nonlinearities modify the conclusions of Section 3.2? We further investigate in this sub-section whether asymmetries exist in the estimates of the Okun law within the identified regimes. We use threshold models to detect nonlinearities and asymmetries in models that we otherwise treat as linear. See Hansen (1997), Tong and Lim (1980), Tong (1983, 1990, and 2001) for a broad treatment of the different classes of threshold autoregressive (TAR) models.

Estimation of TAR models requires that we determine the threshold variable. The choice of the threshold variable, however, is not straightforward, since the underlying economic theory gives no clues as to the identity of the threshold variable. In the threshold regression literature, however, researchers usually consider two main specifications, each defined by the identity of the threshold variable. The first is the self-exciting autoregressive threshold model (SETAR), where the threshold variable is the dependent variable with delay parameter *d* (i.e., $q_t = \Delta U_{t-d}$).

This model combines the autoregressive distributed lag model with a lagged threshold dependent variable. The second specification is the conventional threshold model (TAR), where the threshold variable is not the lagged dependent variable. In this case, we employ $q_t = \Delta GDP_{t-d}$ as a threshold variable with delay parameter *d*. To find the threshold variable, we conduct a search over $\{\Delta U_{t-d}, \Delta GDP_{t-d}\}$. That is, we estimate eight models -- four SETAR and four TAR models -- each with delay parameter *d* ranging from 1 to 4 and choose the specification that minimizes the residual sum of squares.[7] As previously mentioned, we employ the Bai and Perron (1998) methodology and not the fixed regressor bootstrap testing proposed by Hansen (1999). Consequently, the discussion of breakpoint testing and estimation generally applies in the current context.

We first consider the existence of a threshold effect. Table 4 presents the Bai and Perron global and sequential tests for the threshold models applied to each of the three regimes. For the 1949Q2-1973Q2 and 1973Q3-2009Q2 regimes, we set the maximum number of thresholds equal to 3, while for the 2009Q3-2015Q4 regime, because of the relative shortness of the series, we set the maximum number of thresholds equal to 1.[8] The search procedure obtains ΔGDP_{t-1} as the threshold parameter in the first regime, ΔGDP_{t-4} in the second regime, and ΔU_{t-3} in the third regime. Thus, we use the TAR model in the first and second regimes, and the SETAR model in the third regime. Such change itself confirms that a structural change occurred.

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[7] The parameter space for the delay parameter is discrete, which implies that the OLS estimate of the delay parameter is super consistent. Thus, in making inferences about the other parameters, we can treat the delay parameter as if it is certain (Barnes, 1999).

[8] The parameter space for the delay parameter is discrete, which implies that the OLS estimate of the delay parameter is super consistent. Thus, in making inferences about the other parameters, we can treat the delay parameter as if it is certain.

The UD max and WD max statistics, as well as the $\sup F_T(k)$ statistics reject the null hypothesis of no threshold. That is, both tests indicate that at least one threshold exists. This finding rejects the linearity hypothesis in favor of the threshold model. The $\sup F_T(\ell+1|\ell)$ statistics, on the other hand, indicate that only one threshold state exists in each of the three regimes.

Table 5 provides the threshold estimates of Okun's relationship in the first regime, 1949Q2-1973Q2. In both states, unemployment rate persistence does not significantly differ from zero. In both states, output growth exerts a significant negative effect on the unemployment rate. Okun's law, however, is stronger when output growth falls below the threshold, and weaker when output growth exceeds or equals the threshold. Below the threshold, the total effect of output growth on the unemployment rate is -0.654, while above the threshold, the total effect is only -0.251. This implies that when output growth in the previous period switches from below to above the threshold, the total short-run effect of output growth falls by more than half. The test of equality of the total effect of output growth across the first and second state rejects the null at the 1-percent level.

Table 6 provides the threshold estimates of Okun's relationship in the second regime, 1973Q3-2009Q2. Unemployment rate persistence does not significantly differ from zero in the first state, but does significantly differ at the 1-percent level in the second state. The difference between the two persistence estimates, however, does not significantly differ from zero. In both states, the contemporaneous effect of output growth on the unemployment rate is significant with the correct sign. The first lag of output growth is significant at the 1-percent level in the first state, while in the second state, the second lag of output growth is significant at the 1-percent level. Okun's law, however, is stronger when output growth falls below the threshold, and weaker when output growth exceeds or equals the threshold. When $\Delta GDP_{t-4} < 0.26$, (i.e., below the threshold), the total effect of output on the unemployment rate is -0.647, while when $\Delta GDP_{t-1} \geq 0.39$, the total effect is only -0.313. Thus, when output growth at lag 4 switches from below to above the threshold, the total short-run effect of output growth falls by more than one third. The test of equality of the total effect of output growth across the first and second state, however, rejects the null only at the 10 percent level. Thus, the main difference between the first and second states in the second regime is limited to the contemporaneous effect of output growth on the unemployment rate.

Table 7 provides the threshold estimates of Okun's relationship in the third regime, 2009Q3-2015Q4. The threshold model shifts from a TAR to a SETAR model. The threshold variable is ΔU_{t-3} and the threshold value is -0.10.[9] This result confirms that important changes have occurred in Okun's law since the end of the Great Recession. A main difference between the two states is the estimate of ΔU_{t-1}. In the first threshold state, when $\Delta U_{t-3} <$ -0.1, the estimate of the coefficient of ΔU_{t-1} is negative and significant at the 5-percent, while in the second threshold state, the estimate is positive and significant at the 5-percent level. The convergence pattern oscillates in the first state, $\Delta U_{t-3} <$ -0.1, while follows a monotonic adjustment in the second, $\Delta U_{t-3} \geq -0.1$. In the first threshold state, only the estimate of the coefficient of the second lag of ΔGDP_t is significantly different from zero at the 5-percent level with the correct sign. In the second threshold state, on the other hand, only the estimate of the

[9] Hansen (1999, 2000) argues that using the lagged endogenous regressors in the threshold model, which are themselves subject to structural change, violates the assumptions for the sup F(k). This, however, may not cause a problem in our case for two reasons. First, each threshold model is embedded in regimes that do not reveal a structural break. Second, while ΔU_t is subject to a structural break, the date of the break occurs in 2009Q3, which coincides, as noted above, with the break detected in the autoregressive distributed lag specification.

first lag of ΔGDP_t is significant with the correct sign. Thus, in the post-Great Recession regime, no contemporaneous effect of output growth on the unemployment rate occurs, but only a short-run delayed effect exists. The cumulative effect is significant at the 1-percent level in both states and is approximately the same in both states. The long-run effects, however, differ significantly.

4. Conclusions

This paper revisits Okun's law for the United States using data over the period 1949Q4-2015Q4. This period is interesting because, among other things, it includes the Great Recession and its aftermath, which, arguably, saw the worst economic downturn since the Great Depression. We consider the "dynamic" version of Okun's law, which we specify as an autoregressive distributed lag model and draw attention to the findings of structural breaks and threshold nonlinearity in interpreting Okun’s law. We find that over the entire sample period, Okun's law experienced two structural breaks, one in 1973, which coincides with the first crude oil shock, and one in 2009, which coincides with the end of the Great Recession. From the econometric analysis, we can conclude that the reaction of the unemployment rate to changes in real GDP differs substantially between the regimes considered. In particular, the OLS estimates for the third regime appear to indicate that Okun's estimates have become insignificant, casting doubts on the effective validity of the relationship. These results assume, however, that the relation between output growth and the unemployment rate is linear. The threshold analysis conducted inside each regime identified by the Bai-Perron methodology rejects the linearity hypothesis in favor of threshold asymmetry. Taken together, the U.S. data confirm that the reaction of the unemployment rate to output growth depends on both the regime and the threshold. When we incorporate these dependencies, we find that Okun's law remains valid in the sense of statistical significance of the estimates of the total effect of output growth on the unemployment rate, despite the lack of statistical significance of the contemporaneous effect. From such perspective, we believe that Okun's law has withstood the empirical challenge of the post-Great Recession period.

Our finding that Okun’s law experiences structural breaks and threshold switching challenges the linearity and stability of this widely believed empirical regularity amongst macroeconomists. In this regard, future research may consider what factors, in particular, caused the structural shift and the non-monotonity observed in the data in the post-Great Recession period. What we know now is that several circumstances have converged to affect the labor markets. At the secular level, the leading edge of the baby boomers began retiring and this significantly contributed to the decline in the rate of labor force participation. At the cyclical level, the financial crisis affected investment and the lack of investment lowered the growth of total factor productivity. These distorsions ultimately suggest that the temporal connection between output and the unemployment rate may no longer conform to a monotonic relationship. Consequently, the predictions and policy implications of Okun’s law may no longer hold in a straightforward way.

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Figure 1: Okun's Law 1949 to 1960

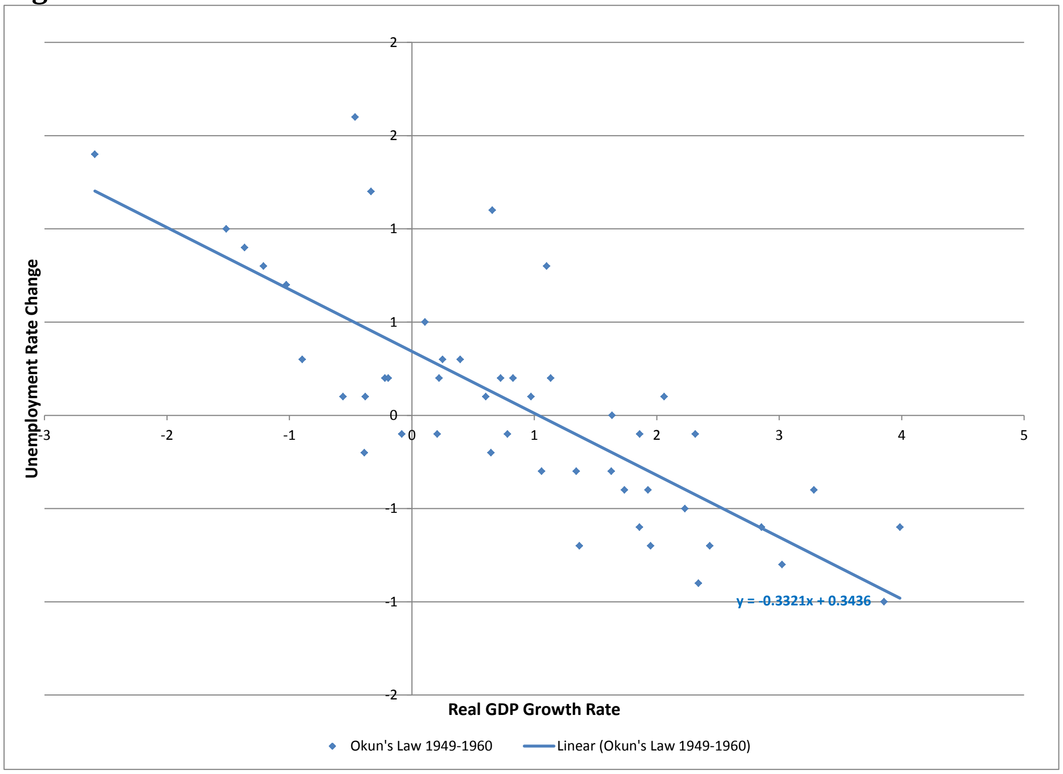

Figure 2: Okun's Law 1949 to 2015

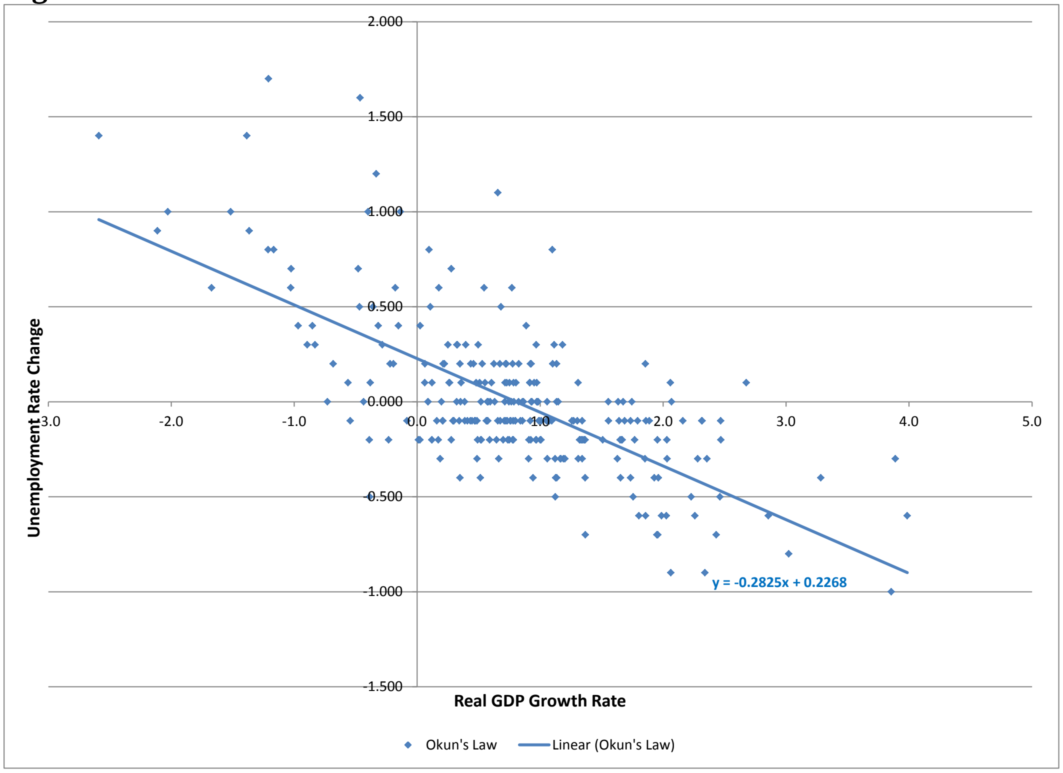

Figure 3: Okun's Law 1949 to 2006, 2007 to 2009, and 2010 to 2015

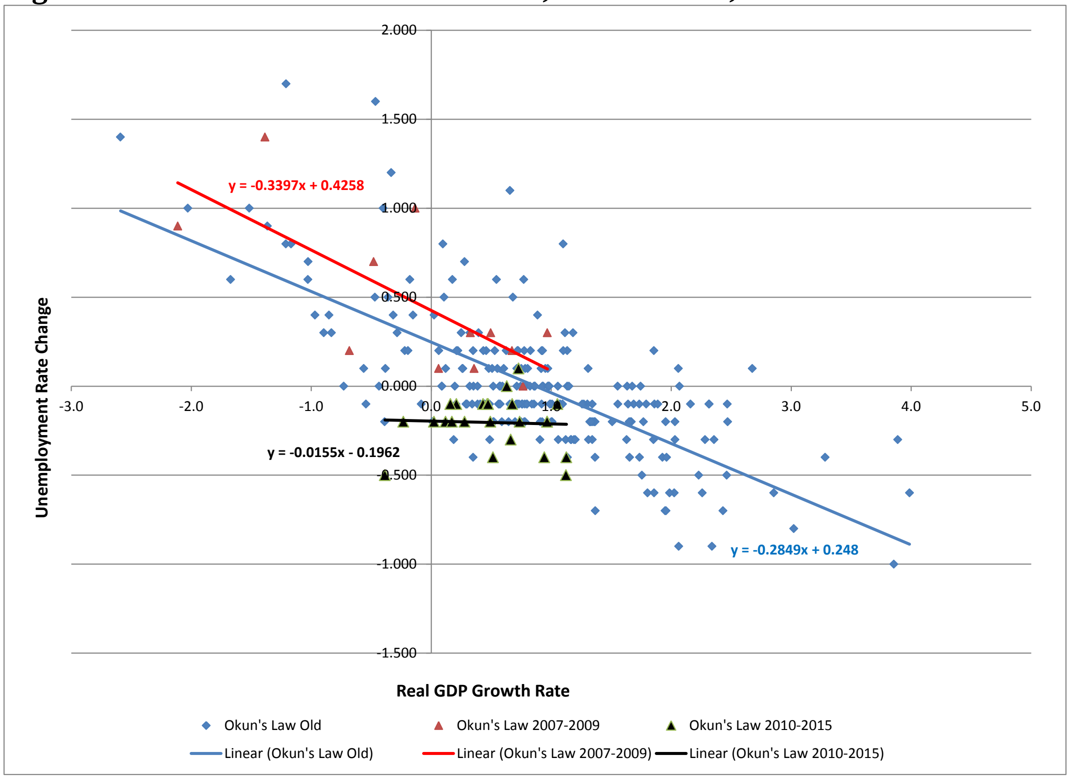

Table 1. Constant parameter estimates of the dynamic linear model: 1949Q2-2015Q4

Variable	Estimate		
Intercept	0.228***		
	(0.038)		
ΔGDP_t	-0.204***		
	(0.021)		
ΔGDP_{t-1}	-0.066***		
	(0.023)		
ΔGDP_{t-2}	-0.024		
	(0.017)		
ΔU_{t-1}	0.315***		
	(0.059)		
Adj.R-squared	0.657	Log likelihood	18.468
Q(4)	1.120	p-value	[0.891]
Q(8)	13.633	p-value	[0.092]
Q(12)	23.362	p-value	[0.025]

Notes. HAC standard errors in parentheses. p-values in brackets.
* Significant at the 10 percent level
** Significant at the 5 percent level
*** Significant at the 1 percent level

Table 2. Bai-Perron specification test results for structural breaks: dynamic version

		Test Statistic	C.V.
UD Max**		142.388	[18.68]
WD Max**		142.388	[20.30]
sup F(1)**		142.388	[18.68]
sup F(2)**		74.794	[16.50]
sup F(3)**		82.606	[15.07]
sup F(1/0)**		142.388	[18.68]
sup F(2/1)**		24.178	[20.57]
sup F(3/2)		18.598	[21.60]
Max No. of breaks	5		
Number of breaks	2		

Notes. Bai-Perron (2003) 5 percent critical values in brackets. The $\sup F_T(\ell)$ is the scaled F statistics from the Bai and Perron (1998) test of ℓ globally optimized breaks against the null of no structural break. The $\sup F_T(\ell+1|\ell)$ is the scaled F statistics from the Bai and Perron (1998) test of ℓ breaks versus $\ell+1$ breaks.

** Significant at the 5-percent level

Table 3. Bai-Perron estimates of the structural break regimes

Variable	1949Q2-1973Q2	1973Q3-2009Q2	2009Q3-2015Q4
Intercept	0.386***	0.281***	-0.053
	(0.071)	(0.055)	(0.052)
ΔGDP_t	-0.220***	-0.225***	0.004
	(0.033)	(0.041)	(0.115)
ΔGDP_{t-1}	-0.127***	-0.071**	-0.076
	(0.029)	(0.035)	(0.059)
ΔGDP_{t-2}	-0.037*	-0.079***	-0.150
	(0.020)	(0.019)	(0.093)
ΔU_{t-1}	0.133**	0.232***	0.089
	(0.066)	(0.085)	(0.159)
Adj.R-squared	0.704	Log likelihood	43.620
Q(4)	6.121	p-value	[0.190]
Q(8)	13.637	p-value	[0.092]
Q(12)	22.309	p-value	[0.034]

Notes. HAC standard errors in parentheses. *p*-values in brackets.
* Significant at the 10 percent level
** Significant at the 5 percent level
*** Significant at the 1 percent level

Table 4. Bai-Perron specification test results for thresholds

	1949Q2-1973Q2		1973Q3-2009Q2		1973Q3-2009Q2	
	Statistic	5% C.V.	Statistic	5% C.V.	Statistic	5% C.V.
UD Max	36.257	17.76	39.716	17.76	79.571	17.14
WD Max	43.588	19.11	47.746	19.11	99.799	18.11
sup F(1)*	33.461	17.66	35.648	17.66	39.609	17.12
sup F(2)*	36.257	14.69	39.716	14.69		
sup F(1/0)*	33.461	17.66	35.648	17.66	39.609	17.12
sup F(2/1)	13.137	19.50	10.128	19.50		
Max No. of thresholds		2		2		1
Number of thresholds		1		1		1
Threshold variable		ΔGDP_{t-1}		ΔGDP_{t-4}		ΔU_{t-3}
Threshold value		0.39		0.26		-0.1

Table 5. Estimates of the TAR threshold model: 1949Q2-1973Q2 regime

	$\Delta GDP_{t-1} < 0.39$		$\Delta GDP_{t-1} \geq 0.39$
Constant	0.559***		0.159**
	(0.062)		(0.068)
ΔGDP_t	-0.272***		-0.201***
	(0.021)		(0.028)
ΔGDP_{t-1}	-0.251**		-0.026
	(0.102)		(0.031)
ΔGDP_{t-2}	-0.131***		-0.023
	(0.046)		(0.025)
ΔU_{t-1}	-0.005		0.106
	(0.171)		(0.071)
Adj.R-squared	0.814	Log Likelihood	22.986
Q(4)	2.532	p-value	[0.639]
Q(8)	4.893	p-value	[0.769]
Q(12)	11.718	p-value	[0.469]
No. Obs.	26		70

Notes. HAC standard errors in parentheses. *p*-values in brackets.
* Significant at the 10 percent level
** Significant at the 5 percent level
*** Significant at the 1 percent level

Table 6. Estimates of the TAR threshold model: 1973Q3-2009Q2 regime

	$\Delta GDP_{t-4} < 0.26$		$\Delta GDP_{t-4} \geq 0.26$
Constant	0.325***		0.237***
	(0.079)		(0.044)
ΔGDP_t	-0.315***		-0.165***
	(0.041)		(0.026)
ΔGDP_{t-1}	-0.173***		-0.031
	(0.048)		(0.026)
ΔGDP_{t-2}	0.021		-0.115***
	(0.029)		(0.029)
ΔU_{t-1}	0.301		0.256***
	(0.251)		(0.055)
Adj.R-square	0.751	Log likelihood	44.724
Q(4)	2.189	p-value	[0.701]
Q(8)	6.803	p-value	[0.558]
Q(12)	11.718	p-value	[0.469]
No. Obs	28		116

Notes. HAC standard errors in parentheses. *p*-values in brackets.
* Significant at the 10 percent level
** Significant at the 5 percent level
*** Significant at the 1 percent level

Table 7. Estimates of the SETAR threshold model: 2009Q3-2015Q4 regime

	$\Delta U_{t-3} < -0.10$		$\Delta U_{t-3} \geq -0.10$
Constant	-0.219**		0.127***
	(0.084)		(0.156)
ΔGDP_t	-0.071		-0.063
	(0.069)		(0.061)*
ΔGDP_{t-1}	-0.051		-0.275***
	(0.062)		(0.013)
ΔGDP_{t-2}	-0.151**		0.046**
	(0.063)		(0.020)
ΔU_{t-1}	-0.562**		0.282***
	(0.264)		(0.017)
Adj.R-squared	0.548	Log likelihood	21.868
Q(4)	1.208	p-value	[0.877]
Q(8)	8.311	p-value	[0.404]
Q(12)	12.477	p-value	[0.408]
No. Obs	15		11

Notes. HAC standard errors in parentheses. *p*-values in brackets.

* Significant at the 10 percent level

** Significant at the 5 percent level

*** Significant at the 1 percent level