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Keywords: Unemployment rate; Unit root tests; Structural breaks; Half-lives.

JEL Codes: C22; E24.

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Abstract

The 2008 economic downturn in the United States resulted in a wave of contractionary effects across many OECD countries. Because of its severity, the Great Recession may have worsened the pattern of unemployment persistence across these countries, where the unemployment rate has remained higher than its pre-crisis levels. This paper investigates the pattern of the unemployment rate in the United States and other 28 OECD countries before and after the 2008 Great Recession. To detect possible changes in the pattern of unemployment persistence in OECD countries, we employ a mean bias-corrected estimation of the persistence parameter with a rolling window of five years. In addition, we estimate the most likely date of change in the trend function of unemployment to test whether there was any significant change in the pattern of unemployment persistence after the Great Recession. Finally, we propose and apply a bootstrap permutation test to verify whether the magnitude of the half-lives and impulse response functions across OECD countries have changed after the Great Recession.

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1 Introduction

The 2008 economic downturn in the United States resulted in a wave of contractionary effects in production and employment in many countries of the Organisation for Economic Co-operation and Development (OECD). Some commentators have even called this crisis as the "Great Recession," characterized by an increase in unemployment across OECD countries. Recovery has been slow in several OECD countries, where the unemployment rate has remained higher than its pre-crisis levels. Thus, there is a concern that these contractionary effects reflect changes in structural conditions that may have worsened the pattern of unemployment persistence across these countries. Two recent editions of the OECD Employment Outlook, launched in mid-2014 and mid-2015, recognize that the unemployment persistence at high levels resulted in a rise in structural unemployment in many OECD countries (Organisation for Economic Co-operation and Development (OECD), 2014, 2015). In addition, according to the 2016 edition of the OECD Employment Outlook, about three-quarters of OECD countries are still facing a considerable unemployment gap, as measured by an unemployment rate that is 2 percentage points or more above the pre-crisis level (Organisation for Economic Co-operation and Development (OECD), 2016).

From a theoretical viewpoint, there are two main hypotheses relating unemployment to economic development: the *non-accelerating inflation rate of unemployment* (NAIRU) theory, and the theory of unemployment hysteresis. Proposed by Phelps (1967) and Friedman (1968), the natural rate of unemployment theory states that unemployment is a mean-reverting process, converging to a natural rate in the long-run, the *non-accelerating inflation rate of unemployment* (NAIRU) hypothesis. However, Blanchard and Summers (1987) proposed the unemployment hysteresis hypothesis, arguing that economic fluctuations have permanent effects on the level of unemployment due to labor market restrictions.

These two hypotheses are verifiable by applying unit root tests on unemployment rates. Under the null hypothesis of unemployment hysteresis, the level of unemployment follows a unit root process. Thus, rejecting a unit root provides evidence for the natural rate hypothesis, but finding evidence of a unit root supports the hysteresis hypothesis. Elmskov and MacFarlan (1993), Mitchell (1993), Røed (1996), Røed (2002), and Camarero and Tamarit

(2004) suggested that, in general, unemployment rates exhibit hysteresis in OECD countries, by applying conventional unit root tests such as the augmented Dickey-Fuller (ADF) test. On the other hand, Røed (1997), Song and Wu (1998), Papell, Murray, and Ghiblawi (2000), Camarero, Carrion-i-Silvestre, and Tamarit (2006), Lee and Chang (2008), and Lee, Lee, and Chang (2009), among others, support the stationarity on the unemployment rates in OECD countries, by employing different methods from the standard ADF test, such as panel unit root tests, the Zivot and Andrews test (Zivot and Andrews, 1992), and unit root tests with endogenous break points. These mixed findings arise because the ADF test displays low power when the span of the data is not long enough or when a structural break is ignored (Perron, 1989).

Because of its severity, the Great Recession may have generated a structural break in the time path of the unemployment rate across OECD countries. Thus, this paper aims to investigate the pattern of the unemployment rate in the United States and other 28 OECD countries before and after the 2008 Great Recession. As in Cuestas, Gil-Alana, and Staehr (2011), we consider the Great Recession as a big event. Following Banerjee, Lumsdaine, and Stock (1992) and Zivot and Andrews (1992), we assume that big events and external shocks affects the economic activity in a permanent or transitory way depending on the nature and on the magnitude of the persistence of key macroeconomic variables, such as the unemployment rate.

To detect possible changes in the pattern of unemployment persistence in OECD countries, we employ mean bias-corrected parameter estimation with a rolling window of five years. We investigate patterns of unemployment persistence before and after the Great Recession using monthly unemployment data for 29 OECD countries and 4 different groupings of countries.

We estimate the most likely change date in the trend function of unemployment by computing an unbiased scalar measure of persistence; this allows us to test whether there is a significant change in the unemployment persistence after the Great Recession. This unbiased scalar measure of persistence yields more accurate and useful information for stabilization and especially structural policies in both the cumulative impulse response and half-life dimensions. Cheng, Durmaz, Kim, and Stern (2012), among others, recognize that the statistical conclusions regarding the common component of state unemployment rates appears to be heavily dependent on the inclusion of most recent data. The null of nonstationarity (hysteresis) is easily rejected using data up through the end of last expansion; however, nonstationarity is easily accepted if the data after the Great Recession are included. Recent events are crucial to our understanding of unemployment. To our knowledge, we are unaware of any other empirical work that computes an unbiased scalar measure of unemployment persistence in OECD countries taking into account the influence of big events like the Great Recession.

Cuestas et al. (2011) investigate the behavior of the unemployment rate in eight countries of Central and Eastern Europe using monthly data from January 1998 to December 2007. They adopt many different approaches to study the nature of unemployment persistence, including fractional integration analysis, unit root with or without structural breaks, and impulse response functions. When structural breaks are considered, only Lithuania has an unemployment rate following a mean-reverting process. Meanwhile, when a fractional integration approach is used, the results of the ARFIMA(1,d,0) show that unemployment persistence is high in all countries.

Jiménez-Rodríguez and Russo (2012) apply a vector autoregression (VAR) model to investigate the behavior of unemployment persistence in five OECD countries that pursued a partial labor market reform program during the 1990's, namely, Italy, Germany France, Spain and the United Kingdom. They find that partial labor market reforms have increased significantly the employment responsiveness to output shocks. Cheng et al. (2012) examine the pattern of unemployment persistence in the U.S. by using data which extend over the Great Recession. They investigate the nature (mean reverting vs. non mean reverting) of the unemployment rate for all the U.S. states from 1976 Q1 to 2010 Q2 by considering the recent labor market turmoil generated by the recent crisis as a truly national shock. They consider the presence of cross-section dependence between the U.S. states and obtain estimates of the half-life and rolling window procedure that are absent in previous studies. When more recent data are included and cross-section dependence is considered, they find strong evidence of hysteresis, with a half-life of six to fourteen years. Fosten and Ghoshray (2011) also adopt a mean-reverting vs. non mean-reverting analysis and offer a comprehensive review of the unit root tests applied to the unemployment rate in OECD countries. However, their results are not comparable with ours because they use lower frequency, annual data on unemployment, which tend to smooth the autoregressive coefficient.

Therefore, this paper contributes to the literature on unemployment persistence in OECD countries in the following ways. First, we estimate a bootstrap mean-unbiased scalar measure of unemployment persistence for most of the OECD countries, which is absent in other studies that are typically focused on the dichotomy between mean-reverting vs. unit root. For comparison purposes, we also present the results of a stationary/nonstationary approach. Second, we apply a rolling window procedure of five years to infer the pattern of changes in our measure of unemployment persistence over the years for each country of our sample. This approach detects changes in the unemployment persistence over time, in the presence of big events such as the Great Recession. Finally, by considering the Great Recession as a big event, we propose and apply a bootstrap permutation test to verify whether the half-lives and impulse response functions in OECD countries have changed significantly after the Great Recession. We find evidence that there is a statistically significant decrease in unemployment persistence across OECD countries after the Great Recession.

We follow the approach of Pivetta and Reis (2007) and Kim (2003) to measure the persistence of macroeconomic variables, but take several steps further. First, we apply the Perron and Rodríguez (2003)'s procedure to make inferences about the general notion of stationary/nonstationary time series, gathering information about the most probable date of change in the trend path of unemployment for partitions of the sample. Second, we estimate a rolling window of five years to follow the path of the unemployment persistence over time. Finally, we employ a bootstrap permutation test to investigate whether such crisis has changed the magnitude of half-lives and impulse response functions across OECD countries. As it turns out, we reject the null hypothesis of no change in the half-lives and twelve-month impulse response functions at the 10% significance level.

The remainder of the paper is organized as follows. Section 2 describes the unit root tests and the econometric methodology. Section 3 presents and discusses the main empirical results, whereas Section 4 reports the main conclusions.

2 Econometric Methodology

2.1 Unit root tests

Since the influential paper of Perron (1989), macroeconomic time series have been best construed as stationary fluctuations around a deterministic trend function if allowance is made for the trend function to exhibit occasional changes. However, Christiano (1992) criticized this postulate by arguing that the choice of the break points had to be viewed as correlated with the data. Subsequently, Zivot and Andrews (1992), Banerjee et al. (1992), and Perron (1997) proposed unit root tests with unknown break points. In this paper, following the approach of Perron and Rodríguez (2003), we treat potential break points as occurring at unknown dates and use the M-tests, proposed by Stock (1999) and further analyzed by Perron and Ng (1996). The M-tests have smaller size distortions than other unit root tests have when the errors present strong negative serial correlation. In addition, the method of Perron and Rodríguez (2003) uses local to unity generalized least squares (GLS) detrending of the data that allows substantial gains in power.

Employing a model common in the previous literature, we assume that the unemployment rate u_t, at time $t \in \{0, 1, \ldots, T\}$, was generated under the null hypothesis of hysteresis as

$$u_t = d_t + \varepsilon_t, \ t = 0, \ldots, T, \tag{1}$$

$$\varepsilon_t = \alpha_1 \varepsilon_{t-1} + \eta_t, \tag{2}$$

where $d_t = \psi' z_t$, for a set of deterministic components z_t, and η_t is an unobserved stationary zero-mean process such that $\eta_t = \sum_{j=0}^{\infty} \gamma_j v_{t-j}$ with $\sum_{j=0}^{\infty} j|\gamma_j| < \infty$, where $\{v_t\}$ is a martingale difference sequence. We assume $\varepsilon_0 = 0$, but inference can be done under a weaker moment requirement that $E(\varepsilon_0^2) < \infty$ (see Perron and Rodríguez, 2003). For any unemployment rate series u_t, with deterministic components z_t, we defined the transformed series $u_t^{\bar{\alpha}}$

and $z_t^{\bar{\alpha}}$ by

$$u_t^{\bar{\alpha}} = \left(u_0, \left(1 - \bar{\alpha}L\right) u_t\right), \ t = 0, \ldots, T, \tag{3}$$

$$z_t^{\bar{\alpha}} = \left(z_0, \left(1 - \bar{\alpha}L\right) z_t\right), \ t = 0, \ldots, T, \tag{4}$$

where L is the lag operator. Let $\hat{\psi}$ be the estimate that minimizes

$$S^* \left(\psi, \bar{\alpha}, \delta\right) = \sum_{t=0}^{T} \left(u_t^{\bar{\alpha}} - \psi' z_t^{\bar{\alpha}}\right)^2, \tag{5}$$

and we denote the minimized value of $S^* \left(\psi, \bar{\alpha}, \delta\right)$ by $S(\bar{\alpha}, \delta)$, with a parameter $\delta \in (0, 1)$ defining the date of the structural change in the series. In this paper, we test the null hypothesis of hysteresis through a model that allows for a structural change in the slope of z_t (Model I) and a model that allows for a structural change in both the intercept and the slope of z_t (Model II).

In Model I, allowing for a structural change in the slope of z_t, the set of deterministic components of z_t in (1) is given by $z_t = \{1, t, 1\left(t > T_C\right)\left(t - T_C\right)\}$, where $1(A)$ is the indicator function of the event A and T_C is the time of the structural change. For simplicity, we assume that $T_C = T\delta$, for some $\delta \in (0, 1)$. Hence, the deterministic component is given by

$$d_t = \mu_1 + \beta_1 t + \beta_2 \left(t - T_C\right) 1\left(t > T_C\right), \tag{6}$$

where $\hat{\psi}(\delta) = (\hat{\mu}_1, \hat{\beta}_1, \hat{\beta}_2)'$ is the vector of estimates that minimizes (5). This specification is similar to the 'additive outlier model' of Perron (1989), where a change in the slope is allowed but both segments of the trend function are joined at the time of the break. In Model II, allowing for a structural change in the intercept and slope of z_t, we have $z_t = \{1, 1\left(t > T_C\right), t, 1\left(t > T_C\right)\left(t - T_C\right)\}$. Then, the deterministic component of Model II is

$$d_t = \mu_1 + \mu_2 1\left(t > T_C\right) + \beta_1 t + \beta_2 \left(t - T_C\right) 1\left(t > T_C\right), \tag{7}$$

where $\hat{\psi}(\delta) = (\hat{\mu}_1, \hat{\mu}_2, \hat{\beta}_1, \hat{\beta}_2)'$ is the vector of estimates that minimizes (5).

The unit root test of Perron and Rodríguez (2003) is performed using the following two-step procedure. First, the unemployment rate series is detrended using

$$\tilde{u}_t = u_t - \hat{\psi}' z_t, \tag{8}$$

where $\hat{\psi}$ minimizes (5). The test is then implemented using the t-statistic for $b_0 = 1$ in the regression:

$$\Delta \tilde{u}_t = b_0 \tilde{u}_{t-1} + \sum_{j=1}^{k} b_j \Delta \tilde{u}_{t-j} + e_{tk}, \tag{9}$$

where $\Delta = (1 - L)$ is the difference operator. We denote this test by either $ADF^{GLS,I}(\delta)$ or $ADF^{GLS,II}(\delta)$ depending on using Model I or II respectively.

To select the break date, we follow the approach of Perron (1997) and Perron and Rodríguez (2003), and choose the break point such that the absolute value of the t-statistic on the coefficient of the change in the slope is maximized. For example, in Model I of equation (6), let $\hat{\beta}_2(\delta)$ be the GLS estimate of β_2 and $t_{\hat{\beta}_2}(\delta)$ be its associated t-statistic. Then, the break date is selected using

$$\hat{\delta} = \underset{\delta \in [\epsilon, 1-\epsilon]}{\arg\max} \left| t_{\hat{\beta}_2}(\delta) \right|, \tag{10}$$

where ϵ is some small number. We follow Perron and Rodríguez (2003) and choose $\epsilon = 0.15$ throughout this paper.

As our test requires the estimation of the augmented regression (9), we employ a data-dependent method to select the optimal order k of the autoregression in equation (9). Data-dependent methods lead to test statistics having better properties than if a fixed k is chosen a priori, see Hall (1994) and Ng and Perron (1995). We confine the search for the best value of k in a range $[0, k_{\max}]$ and estimate all regressions using the same number of effective observations, $T^* = T - k_{\max}$. To select the optimal lag order k, we follow Ng and Perron

(2001) and use the Modified Akaike Information Criterion (MAIC) defined by

$$k^* = \operatorname{argmin}_{k \in [0, k_{\max}]} \left\{ \log \left(s_{ek}^2 \right) + \left(2 \left(\hat{\tau}_T(k) + k \right) \right) / T^* \right\}, \tag{11}$$

with $s_{ek}^2 = T^{*-1} \sum_{t=k_{\max}+1}^{T} \hat{e}_{tk}^2$ and $\hat{\tau}_T(k) = (s_{ek}^2)^{-1} \hat{b}_0^2 \sum_{t=k_{\max}+1}^{T} \tilde{u}_{t-1}^2$, with $\hat{e}_{tk}$ and $\hat{b}_0$ obtained from the augmented regression (9) estimated from $t = k_{\max} + 1$ to T. Ng and Perron (2001) show that the MAIC works as well as standard information criteria when the extent of correlation is mild but provides unit root tests having better size properties with a negative MA component.

The unit root test of Perron and Rodríguez (2003) ignores the case where only a change in the intercept is allowed, since it is a special case of what Elliott, Rothenberg, and Stock (1996) consider as a slowly evolving deterministic component. Thus, to find the most likely structural break in the path of each unemployment rate series, we estimate Model A of Perron (1997) and test the significance of the structural break dummy. Model A of Perron (1997) allows only for a change in the intercept under both the null and alternative hypotheses. Besides, this change is assumed to happen gradually and in a way depending on the correlation structure of the noise function. Then, the unit-root test of Perron (1997) is performed using the t-statistic for testing $\alpha = 1$ in the following OLS regression:

$$u_t = \mu_1 + \mu_2 DU_t + \beta_1 t + \delta D\left(T_C\right)_t + \alpha u_{t-1} + \sum_{j=1}^{k} c_j \Delta u_{t-j} + e_t, \tag{12}$$

where $DU_t = 1\left(t > T_C\right)$ and $D\left(T_C\right)_t = 1\left(t = T_C + 1\right)$. To select the optimal order k of the autoregression in equation (12), we apply a general to specific recursive procedure based on the t-statistic associated with the last lag in the estimated autoregression, the t-sig method of Perron (1997). This procedure selects the value of k, say k^*, such that the coefficient on the last lag in an autoregression of order k^* is significant and the last coefficient in an autoregression of order greater than k^* is insignificant, up to a maximum lag order of 12. We select the break in the same way as in the unit root test of Perron and Rodríguez (2003), maximizing the absolute value of the t-statistic of μ_2 in (12) as described in equation (10).

2.2 A scalar measure of unemployment persistence

In this paper, we define persistence as the long-run effect of a shock to the unemployment level u_t. To estimate the unemployment persistence level, we follow the approach of Kim (2003) that delivers better finite sample properties when the auto-regressive coefficient of u_t is close or equal to one. The algorithm proposed by Kim (2003) presents two advantages over the existing procedures. First, its bootstrap mean bias-corrected estimator is simulated conditional on the last p observations in the sample path, which incorporates the conditionality of AR forecasts into bootstrap bias estimation. This is an attractive property since AR forecasts are conditional on the last p observations. Second, this estimator corrects for biases in the least-squares estimators for all parameters in the model simultaneously. This is different from the bias-corrected estimators of Andrews and Chen (1994) and Roy and Fuller (2001) that correct for the bias in the least-squares estimator for the persistence parameter first, and then re-estimate the other parameters of the model given the bias-corrected estimate of the persistence parameter.

We specify the unemployment series as an $AR(p)$ model of the following form:

$$\begin{aligned} u_t &= \mu' + \beta' t + \varepsilon_t, \\ \varepsilon_t &= \alpha_1 \varepsilon_{t-1} + \ldots + \alpha_p \varepsilon_{t-p} + \eta_t, \end{aligned} \tag{13}$$

where $\eta_t \sim iid(0, \sigma^2)$. We can rewrite the model (13) as follows:

$$u_t = \mu + \beta t + \alpha_1 u_{t-1} + \ldots + \alpha_p u_{t-p} + \eta_t, \tag{14}$$

where $\mu = \mu'(1 - \alpha_1 - \ldots - \alpha_p) + (\alpha_1 + 2\alpha_2 + \ldots + p\alpha_p)\beta'$ and $\beta = \beta'(1 - \alpha_1 - \ldots - \alpha_p)$. Rewriting equation (14), we have:

$$u_t = \mu + \beta t + \alpha u_{t-1} + \iota_1 \Delta u_{t-1} + \ldots + \iota_{p-1} \Delta u_{t-p+1} + \eta_t, \tag{15}$$

where α is the persistence parameter that measures the degree of persistence of the AR model. We can relate the parameters of equation (14) and (15) as follows:

$$\begin{aligned}
&\alpha = \alpha_1 + \ldots + \alpha_p, \\
&\alpha_1 = \alpha + \iota_1, \alpha_j = -\iota_{j-1} + \iota_j, \text{ for } 2 \leq j \leq p-1, \\
&\alpha_j = -\iota_{j-1}, \text{ for } j = p.
\end{aligned} \tag{16}$$

Given the observed unemployment series $\{u_t\}_{t=1}^T$, we denote the least-squares estimator for $\gamma = (\mu, \beta, \alpha)$ by $\hat{\gamma} = (\hat{\mu}, \hat{\beta}, \hat{\alpha})$, where $\hat{\alpha} = \hat{\alpha}_1 + \ldots + \hat{\alpha}_p$. The resulting residuals are defined as $\{\hat{\eta}_t\}_{t=p+1}^T$. Since the least-squares estimator of α is biased, we apply the bootstrap mean bias-corrected estimator of Kim (2003) as follows:

1. Regress model (14) by least-squares and get the residuals $\{\hat{\eta}_t\}_{t=p+1}^T$. Then generate a pseudo data set based on the backward AR(p) model as

$$u_t^* = \hat{\mu} + \hat{\beta}t + \hat{\alpha}_1 u_{t+1}^* + \ldots + \hat{\alpha}_p u_{t+p}^* + \eta_t^*, \tag{17}$$

   where $\{u_t\}_{t=T-p+1}^T$ are the starting values and η_t^* is a random draw with replacement from $\{\hat{\eta}_t\}_{t=p+1}^T$.

2. Obtain the bootstrap parameter estimate for γ, $\gamma^* = (\mu^*, \beta^*, \alpha^*)$, by regressing u_t^* on $(1, t, u_{t+1}^*, u_{t+p}^*)$;

3. Repeat steps 1 and 2 B times to generate B sets of bootstrap parameter estimates for γ, $\{\gamma^*(j)\}_{j=1}^B$;

4. Calculate the bias of $\hat{\gamma}$ as $Bias(\hat{\gamma}) = \bar{\gamma^*} - \hat{\gamma}$, where $\bar{\gamma^*} = \sum_{j=1}^B \gamma^*(j)/B$;

5. Obtain the bias-corrected estimator for γ as $\hat{\gamma}^c = \hat{\gamma} - Bias(\hat{\gamma})$.

The bootstrap mean bias-corrected estimator of the persistence parameter, $\hat{\alpha}^c$, correctly provides the length of time until the impulse response function of a unit shock to unemployment is equal to half of its original magnitude - the half-life of a unit shock. This measure characterizes the likely duration of unemployment in all countries of the sample and is defined as follows:

$$HL := HL(\alpha) = \frac{\log(1/2)}{\log(\alpha)}. \tag{18}$$

It is well known that the least-squares estimator of the slope coefficient, in the AR(1) model with linear time trend, is biased. However, Andrews (1993) suggests that this bias seems to be absent when we specify a pure AR(1) model without either drift or linear trend. We specify an AR(1) model with drift and time trend and apply Kim (2003)'s procedure to estimate all coefficients displayed in Figure 1, obtained by rolling window regressions. This procedure allows us to see the evolution of the unemployment persistence over time and the effect of Great Recession.

2.3 A permutation test for differences in unemployment persistence before and after the Great Recession

In this subsection, we propose a bootstrap permutation test to verify whether the half-lives and impulse response functions in OECD countries have changed significantly after the Great Recession. This test is discussed in Efron and Tibshirani (1993). The idea is free of mathematical or behavioral assumptions, since the permutation test works only with the empirical distributions of the samples.

Let $\boldsymbol{\alpha}^b = (\alpha_1^b, \ldots, \alpha_m^b)$ and $\boldsymbol{\alpha}^a = (\alpha_1^a, \ldots, \alpha_n^a)$ be random vectors of monthly unemployment persistence parameters before and after the Great Recession respectively, for $m \leq T_C$ and $T_C < n \leq T$, assuming the date of the structural break T_C is known. We assume that $\boldsymbol{\alpha}^b$ and $\boldsymbol{\alpha}^a$ are drawn from two possibly different distributions F and G, before and after the Great Recession respectively. Let $\mathbf{HL}(\boldsymbol{\alpha}^b) = \big(HL(\alpha_1^b), \ldots, HL(\alpha_m^b)\big)$ and $\mathbf{HL}(\boldsymbol{\alpha}^a) = (HL(\alpha_1^a), \ldots, HL(\alpha_n^a))$ be the half-lives implied by $\boldsymbol{\alpha}^b$ and $\boldsymbol{\alpha}^a$, respectively. We want to test the null hypothesis of $H_0 : F = G$. This means that F and G assign equal probabilities to all sets of unemployment persistence parameters: $\Pr_F(A) = \Pr_G(A)$, for any subset A of the common sample space of $\boldsymbol{\alpha}^b$ and $\boldsymbol{\alpha}^a$. Then, under H_0, there is no difference between the probabilistic behavior of the random vectors of half-lives $\mathbf{HL}(\boldsymbol{\alpha}^b)$ and $\mathbf{HL}(\boldsymbol{\alpha}^a)$. For our two-sample problem, we define the difference between the half-lives as:

$$\hat{\theta} = \overline{\mathbf{HL}}(\boldsymbol{\alpha}^a) - \overline{\mathbf{HL}}(\boldsymbol{\alpha}^b), \tag{19}$$

where $\overline{\mathbf{HL}}(\boldsymbol{\alpha}^a) = \sum_{j=1}^{n} HL(\alpha_j^a)/n$ and $\overline{\mathbf{HL}}(\boldsymbol{\alpha}^b) = \sum_{j=1}^{m} HL(\alpha_j^b)/m$. We reject the null hypothesis $H_0 : F = G$ whenever the test statistic $\hat{\theta}$ is significantly different from zero, where $\hat{\theta}$ is the mean-difference between the half-lives of the unemployment persistence after and before the Great Recession. Analogously, we test the null hypothesis of equal twelve-month impulse response functions before and after the Great Recession. Let $\mathbf{IRF}(\boldsymbol{\alpha}^b) = \big(IRF(\alpha_1^b), \ldots, IRF(\alpha_m^b)\big)$ and $\mathbf{IRF}(\boldsymbol{\alpha}^a) = (IRF(\alpha_1^a), \ldots, IRF(\alpha_n^a))$ be the twelve-month impulse response functions implied by $\boldsymbol{\alpha}^b$ and $\boldsymbol{\alpha}^a$, respectively. We want to test the null hypothesis of $H_0 : F = G$. Then, we define the difference between the half-lives as:

$$\hat{\theta}_{IRF} = \overline{\mathbf{IRF}}(\boldsymbol{\alpha}^a) - \overline{\mathbf{IRF}}(\boldsymbol{\alpha}^b), \tag{20}$$

where $\overline{\mathbf{IRF}}(\boldsymbol{\alpha}^a) = \sum_{j=1}^{n} IRF(\alpha_j^a)/n$ and $\overline{\mathbf{IRF}}(\boldsymbol{\alpha}^b) = \sum_{j=1}^{m} IRF(\alpha_j^b)/m$. Given the estimated $\hat{\theta}$ and $\hat{\theta}_{IRF}$, we define the achieved significance value (ASL) of the test as:

$$ASL = \Pr_{H_0}\left(\theta_0 \geq \hat{\theta}\right), \tag{21}$$

$$ASL_{IRF} = \Pr_{H_0}\left(\theta_0 \geq \hat{\theta}_{IRF}\right), \tag{22}$$

where θ_0 is the value of θ under the null hypothesis. Hence, the smaller the values of ASL and ASL_{IRF}, the stronger the evidence against H_0. Since ASL and ASL_{IRF} are a degree of the credibility of H_0, low values of ASL and ASL_{IRF} support the rejection of the null hypothesis. The proposed test combines all the m plus n observations, where $T = m + n$ for all countries in the sample, and takes random samples of unemployment persistence parameters of sizes m and n before and after the Great Recession respectively, without replacement. We compute $\hat{\theta}$ in (19) and $\hat{\theta}_{IRF}$ in (20), and repeat this procedure 10,000 times.

3 Empirical Results

We use country-level seasonally adjusted unemployment rates of 29 OECD countries and 4 groupings of countries, the EUROAREA zone, the EU28, the G7 countries, and the OECD, spanning from Jan.2000 through Oct.2015. We obtain the data from the OECD.Stat. Due to data availability, we excluded from our sample the following OECD countries: Iceland, Israel, Latvia, New Zealand, Switzerland, and Turkey.

We implement the unit root test $ADF^{GLS}(\delta)$ of Perron and Rodríguez (2003) of equations (8)-(9), which allows for a structural break in the slope of the deterministic components of the unemployment rate series, $ADF^{GLS,I}(\delta)$, and in both the intercept and the slope of the deterministic components of the unemployment rate series, $ADF^{GLS,II}(\delta)$. Table 1 shows that the $ADF^{GLS,I}(\delta)$ test does not reject the null hypothesis of unit root for all unemployment series at the 5% significance level. These results are robust if we allow for a structural break in the intercept and in the slope. Therefore, we do not reject the null hypothesis of hysteresis in unemployment in all countries and groupings of countries of the sample at the 5% significance level.

To find the most likely structural break in the path of each unemployment rate series, we estimate Model A of Perron (1997) in equation (12) and test the significance of the dummy intercept coefficient μ_2, as the procedure of Perron and Rodríguez (2003) lacks a significance test for the dummy intercept coefficient. Table 2 displays the estimates of the most likely structural break date, for each one of the unemployment rate series. There is evidence of statistically significant change in the path of unemployment rate series across all OECD countries and groupings of countries. All dummy intercept coefficients are significant at the 5% level; besides, the estimated structural break dates for the majority of the countries are close to Mar.2008, which is the estimated structural break date of the Great Recession in the US economy. This suggests the presence of a synchronized behavior of the unemployment rate series among the countries of the sample.

Now we estimate the unbiased scalar measure of unemployment persistence, half-lives, and impulse response functions before and after the Great Recession. We also test whether they are equal through the permutation test. We expect that there are changes in the half-lives and in the impulse response functions after the Great Recession, which hit most of the countries in our sample. We excluded the United Kingdom and Luxembourg from our sample because their estimated structural break dates leave insufficient data for inference. We consider the estimated crisis date as the estimated structural break date for each one of the countries presented in Table 2.

Table 3 presents estimates of the unemployment persistence, half-lives, and impulse response functions for 12 months. Equation (15) above states that shocks die out at the rate of $(1 - \alpha)$ per period, where α is the unemployment persistence parameter. If $\alpha = 1$, then the shocks will never die out. We calculated the results of Table 3 for one year (12 months). We estimate the unemployment persistence parameter (α) by the bootstrap mean-bias corrected estimator of Kim (2003) defined in equation (15). We consider the minimal and maximal lag orders of 1 and 6, respectively, and use the BIC criterion to select the optimal lag length of the autoregressions. We specify the backward AR(p) model in (15) with drift and trend, performing 1,000 bootstrap replications.

Table 3 reports that the unemployment persistence parameter estimate is close to unity and higher than 0.9 for most of the countries in the sample, before and after their estimated crisis date. These results are in accordance with the $ADF^{GLS}(\delta)$ unit root tests displayed in Table 1, as all unemployment rate series follow a unit root process. The estimated persistence parameters are close to those estimated using only pre-crisis data by Cuestas et al. (2011) for the six Eastern-European countries: Czech Republic, Estonia, Hungary, Poland, Slovakia, and Slovenia. In addition, for the period before the Great Recession, our results are similar to the findings of Cheng et al. (2012), who found strong evidence of hysteresis in the U.S. economy, with a half-life of six to fourteen years. However, for the period after the Great Recession, we obtain different results for the U.S economy as the estimated half-life is around seven months. Thus, our results support the evidence of hysteresis of the unemployment rate across OECD countries after the Great Recession. In addition, while the average of the half-lives of all countries decreased from 212 to 60 months after the Great Recession, there is an increase in the coefficient of variation of the half-lives of 80% compared with its pre-crisis value. Thus, these results indicate that unemployment duration decreased in many countries (e.g., in Japan, Mexico, and Canada) and the persistence level remained high in others (e.g., in Greece, Hungary, and Ireland).

Table 4 presents the p-values of the achieved significance level (ASL) permutation test for changes in the half-lives and the twelve-month impulse response functions described in (21)-(22) across 27 OECD countries and 4 groupings of countries. For comparison purposes, we also apply a t-test for paired samples under the normality assumption. Table 4 shows that there is a significant change in the half-lives of the unemployment across OECD countries after the Great Recession, as we reject the null hypothesis of no change at the 5% level. We also reject the null hypothesis of equal impulse response functions before and after the Great Recession at the 10% significance level.

Now we analyze the rolling window regression results using five years of data, whose figures are displayed on the Appendix. We estimate the unbiased unemployment persistence parameter (α) by the Kim (2003)'s procedure of equation (15), with $p = 1$ in a model including drift and a linear trend. Figure 1 displays the time path of the unemployment persistence (α) coefficient over five years. The first part of Figure 1 supports the evidence that the US Great Recession and the synchronized behavior in these eight countries produced abrupt changes in the unemployment persistence (α) coefficient around 2008 and 2009, except in Norway. The most affected countries of this group are Canada, Belgium, Mexico, Japan, and Sweden, where the low level of unemployment persistence experienced before the Great Recession changed after the crisis. The most affected country is Canada, and likely for the same reason as Mexico, since they are neighboring economies of the USA. In addition, Austria, Japan, Mexico, and Sweden present a negative trend path in the unemployment persistence parameter over the last years of data.

The second part of Figure 1 suggests that these eight countries experienced only a transitory change in the unemployment persistence (α) coefficient. Among this group of countries, only Australia has recently experienced a significant decline in its unemployment persistence. The third part of Figure 1 reveals changes in the unemployment persistence (α) coefficient around the years 2008 and 2009, except for Luxembourg and Portugal. South Korea presented a striking behavior with a high volatile level of persistence, where the response to shocks has been lower and the adjustment faster when compared with the rest of countries. On the other hand, Italy experienced a permanent increase in the unemployment persistence

(α) coefficient after the Great Recession.

The fourth part of Figure 1 supports the evidence of a synchronized behavior in the path of unemployment persistence. All countries of this group had a transitory but relevant change in the unemployment persistence (α) coefficient around 2008 and 2009, where the estimated coefficient α quickly went back to a near-unity value. For the U.S. economy, our results are similar to the findings of Cheng et al. (2012), whose rolling window analysis exhibits a sudden increase in unemployment persistence in the aftermath of the Great Recession.

In sum, our results reveal a significant change in the pattern of unemployment persistence across OECD countries over 2000-2014. At the 10% significance level, there is evidence of a change in the half-lives and impulse response functions across OECD countries after the Great Recession. Nonetheless, there is a great deal of heterogeneity in the evolution of the unemployment persistence coefficient across individual countries.

Table 1. $ADF^{GLS}(\delta)$ unit root test results

Country or region	Optimal lag	$ADF^{GLS,I}(\delta)$ (-3.89)	Optimal lag	$ADF^{GLS,II}(\delta)$ (-3.89)
1. Australia	3	-3.179	3	-3.364
2. Austria	2	-2.216	2	-2.573
3. Belgium	3	-2.134	3	-2.304
4. Canada	0	-1.637	0	-2.008
5. Chile	6	-2.648	4	-2.791
6. Czech Republic	2	-2.176	5	-2.847
7. Denmark	5	-1.941	11	-2.673
8. Estonia	5	-2.251	9	-2.744
9. Finland	2	-2.678	5	-2.952
10. France	1	-2.292	1	-2.298
11. Germany	2	-2.561	3	-2.941
12. Greece	5	-2.312	9	-2.690
13. Hungary	5	-1.831	10	-2.012
14. Ireland	6	-2.243	6	-2.197
15. Italy	7	-2.717	7	-2.937
16. Japan	3	-1.895	3	-1.921
17. South Korea	4	-2.947	4	-3.145
18. Luxembourg	3	-2.615	3	-2.708
19. Mexico	2	-2.561	2	-2.749
20. Netherlands	3	-2.518	5	-2.844
21. Norway	5	-2.163	5	-2.290
22. Poland	5	-2.414	5	-2.477
23. Portugal	2	-2.203	5	-2.354
24. Slovak Republic	1	-1.767	6	-2.303
25. Slovenia	6	-2.010	9	-2.099
26. Spain	2	-2.313	7	-2.445
27. Sweden	3	-1.974	4	-2.126
28. United Kingdom	4	-2.089	4	-2.010
29. United States	5	-2.829	5	-2.807
30. EUROAREA	2	-2.357	3	-2.714
31. EU28	3	-2.301	3	-2.545
32. G7	2	-2.220	5	-2.377
33. OECD	2	-2.456	5	-2.714

This table presents the results of the unit root tests $ADF^{GLS,I}(\delta)$ and $ADF^{GLS,II}(\delta)$ of Perron and Rodríguez (2003) described in equations (8)-(9). $ADF^{GLS,I}(\delta)$ refers to Model I of equation (6) that allows for a structural change in the slope of the deterministic components of the unemployment rate series. $ADF^{GLS,II}(\delta)$ refers to Model II of equation (7) that allows for a structural change in both the intercept and the slope of the deterministic components of the unemployment rate series. To select the break date, we choose the break point such that the absolute value of the t-statistic on the coefficient of the change of the slope is maximized. The optimal lag lengths were selected by minimizing the Modified Akaike Information Criterion (MAIC) of equation (11), suggested by Ng and Perron (2001). Numbers in parentheses represent the 5% asymptotic critical values of the $ADF^{GLS,I}(\delta)$ and $ADF^{GLS,II}(\delta)$ test statistics taken from Perron and Rodríguez (2003, Table 1, p. 11).

Table 2. Change trend date estimates of unemployment rate series

Country or region	T_C	Optimal lag order	t-statistic for $H_0 : \mu_2 = 0$
1. Australia	Jul.2008	12	2.940***
2. Austria	Apr.2007	12	-3.133***
3. Belgium	Apr.2006	9	-2.768***
4. Canada	Sep.2008	12	6.893***
5. Chile	Apr.2007	6	2.294**
6. Czech Republic	Sep.2008	5	3.751***
7. Denmark	Apr.2008	10	5.410***
8. Estonia	Apr.2008	8	5.203***
9. Finland	Jan.2005	5	-2.969***
10. France	Sep.2008	6	2.690***
11. Germany	Apr.2005	2	-3.542***
12. Greece	Jun.2012	11	-3.733***
13. Hungary	Dec.2012	9	-3.985***
14. Ireland	Jan.2008	7	3.859***
15. Italy	Mar.2011	12	3.957***
16. Japan	Oct.2008	12	5.682***
17. South Korea	Jan.2010	7	-2.055**
18. Luxembourg	Nov.2002	11	2.802***
19. Mexico	May.2008	12	3.350***
20. Netherlands	May.2011	12	3.606***
21. Norway	Oct.2005	3	-5.150***
22. Poland	Mar.2005	8	-4.335***
23. Portugal	Dec.2012	10	-3.695***
24. Slovak Republic	Sep.2008	1	5.434***
25. Slovenia	Nov.2008	6	3.962***
26. Spain	Feb.2008	12	4.638***
27. Sweden	Apr.2008	11	3.311***
28. United Kingdom	Jun.2013	12	-3.384***
29. United States	Mar.2008	12	4.090***
30. EUROAREA	Feb.2008	12	3.834***
31. EU28	Feb.2008	12	4.546***
32. G7	Mar.2008	5	4.277***
33. OECD	Mar.2008	12	5.409***

This table presents the estimates of change trend date of the unemployment rate series. The specified model is described in equation (12), and our null hypothesis is $H_0 : \mu_2 = 0$. T_C denotes the time at which the change in the trend function occurs. The best lag length is selected accordingly to the data-dependent procedure described in Perron (1997) as the t-sig method, and the break date is chosen by maximizing the absolute value of the t-statistic on the coefficient of the change of the slope, as described in equation (10). The maximum lag length is 12. The asymptotic critical values are $-5.15(1\%)$, $-4.84(5\%)$, and $-4.59(10\%)$, which were all extracted from Perron (1997, Table 1, p. 362).

Table 3. Persistence estimates, half-lives, and impulse response functions

Countries	$\hat{\alpha}^b$	$\hat{\alpha}^a$	HL^b	HL^a	$IRF(12)^b$	$IRF(12)^a$
1. Australia	0.955	0.880	15.054	5.422	0.575	0.216
2. Austria	0.999	0.943	692.801	11.811	0.988	0.494
3. Belgium	0.932	0.970	9.843	22.757	0.430	0.694
4. Canada	0.976	0.818	28.533	3.450	0.747	0.090
5. Chile	0.999	0.984	692.801	42.974	0.988	0.824
6. Czech Republic	0.999	0.964	692.801	18.905	0.988	0.644
7. Denmark	1.000	0.914	Inf	7.708	1.000	0.340
8. Estonia	0.874	0.963	5.147	18.385	0.199	0.636
9. Finland	0.848	0.979	4.204	32.659	0.138	0.775
10. France	0.983	0.893	40.426	6.125	0.814	0.257
11. Germany	0.934	0.984	10.152	42.974	0.441	0.824
12. Greece	0.715	0.921	2.066	8.423	0.018	0.372
13. Hungary	0.984	0.991	42.974	76.669	0.824	0.897
14. Ireland	0.916	0.983	7.900	40.426	0.349	0.814
15. Italy	0.999	0.994	692.801	115.178	0.988	0.930
16. Japan	0.976	0.665	28.533	1.699	0.747	0.007
17. South Korea	0.920	0.960	8.313	16.980	0.368	0.613
18. Mexico	0.988	0.665	57.415	1.699	0.865	0.007
19. Netherlands	0.993	1.000	98.674	Inf	0.919	1.000
20. Norway	0.795	0.942	3.021	11.601	0.064	0.488
21. Poland	0.999	0.987	692.801	52.972	0.988	0.855
22. Portugal	0.992	0.898	86.296	6.443	0.908	0.275
23. Slovak Republic	0.998	0.971	346.227	23.553	0.976	0.702
24. Slovenia	0.998	1.000	346.227	Inf	0.976	1.000
25. Spain	0.999	0.999	692.801	692.801	0.988	0.988
26. Sweden	0.998	0.790	346.227	2.941	0.976	0.059
27. United States	0.991	0.895	76.669	6.248	0.897	0.264
28. EUROAREA	1.000	0.993	Inf	98.674	1.000	0.919
29. EU28	0.856	0.998	4.458	346.227	0.155	0.976
30. G7	0.998	0.922	346.227	8.535	0.976	0.377
31. OECD	0.992	0.932	86.296	9.843	0.908	0.430

This table presents the estimates of bootstrap mean bias-corrected estimates of unemployment persistence, half-lives, and impulse response functions in OECD countries of Kim (2003) defined in (15). $\hat{\alpha}^b$ and $\hat{\alpha}^a$ denote the estimated autoregressive parameter before and after the estimated crisis date, respectively. HL^b and HL^a denote the half-lives before and after the estimated crisis date, respectively, which are measured in months. $IRF(12)^b$ and $IRF(12)^a$ denote the impulse response functions before and after the estimated crisis date, respectively, for twelve months. Following Kim (2003), we consider the minimal and maximal lag orders of 1 and 6, respectively, and use the BIC criterion to select the best lag length. We specify the backward AR(p) model in (17) with drift and trend, performing 1,000 bootstrap replications. We excluded Luxembourg and the United Kingdom from our sample because there were few data points after the split in the sample into two parts.

Table 4. Permutation test for half-lives and impulse response functions

Type of test	**Null hypothesis**	p**-value**
Permutation test	Equal HL	0.0079***
t-student test (under normality)	Equal HL	0.0097***
Permutation test	Equal $IRF(12)$	0.0998*
Paired t test (under normality)	Equal $IRF(12)$	0.0987*

This table presents the p-values of the achieved significance level (ASL) permutation test for changes in the half-lives and impulse response functions of the unemployment rates described in (21)-(22), across 27 OECD countries and 4 regions of countries, where *, **, and *** denote significance at the 10%, 5% and 1% levels, respectively. HL denotes the half-lives of equation (18) and $IRF(12)$ denote the impulse response functions of twelve months. The ASL p-values of the permutation test were obtained by performing 10,000 bootstrap replications.

4 Conclusions

The Great Recession may have worsened the pattern of unemployment persistence across OECD countries, where the unemployment rate has remained higher than its pre-crisis levels. This paper investigates the pattern of the unemployment rate in the United States and other 28 OECD countries before and after the 2008 Great Recession. By using a variety of econometric methods, we provide a detailed empirical analysis of the evolution of the unemployment persistence across OECD countries over the period 2000-2014. The rolling window analysis shows that there are important changes in the unemployment persistence coefficient for some countries, as the estimated autoregressive coefficients have different values before and after the Great Recession.

The mean bias-corrected estimator of the unemployment persistence coefficient indicates that there is a significant change in the duration of unemployment across OECD countries after the Great Recession. The bootstrap permutation test results suggest that there are changes in the half-lives and impulse response functions after the crisis. We reject the null hypothesis of no change in the half-lives and twelve-month impulse response functions at the 10% significance level.

In addition, we find evidence of an average decrease in unemployment persistence together with greater dispersion across OECD countries after the Great Recession. In sum, we believe that this paper provides new evidence on the patterns of unemployment persistence in OECD countries and contributes to the literature on the determinants of such patterns and, therefore, on the policies that affect them in a welfare-improving way.

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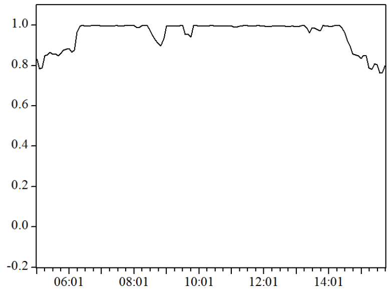
Appendix



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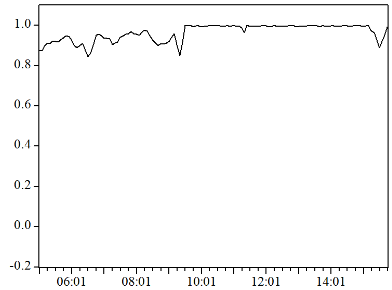

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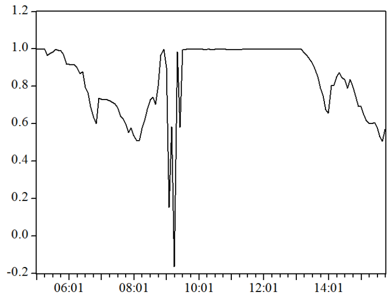

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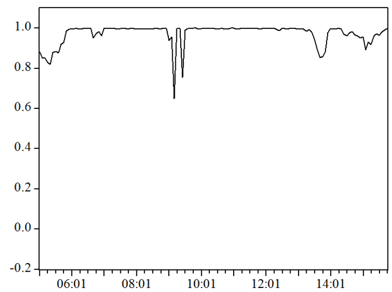

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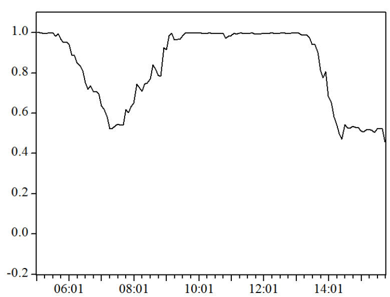

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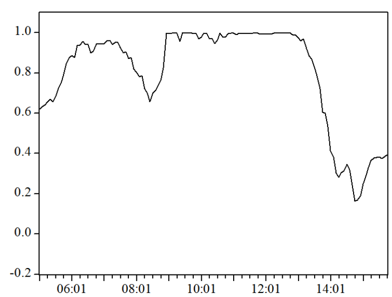

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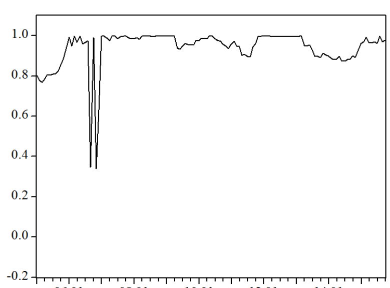

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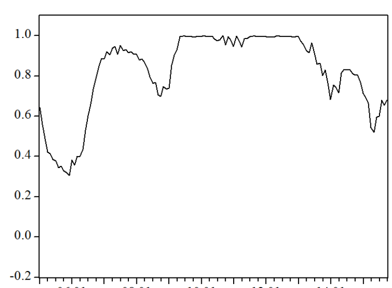
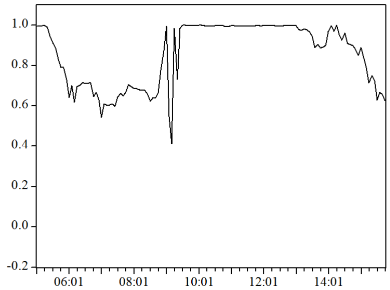
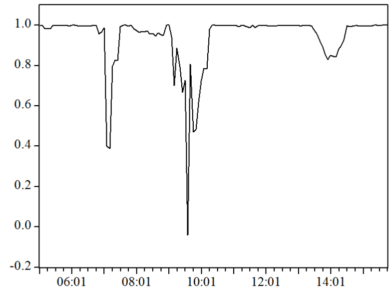


(8) Sweden

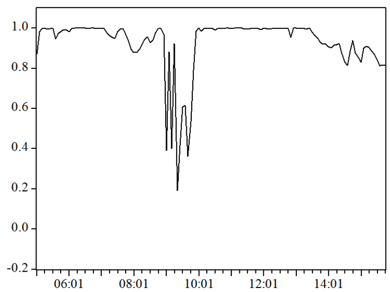
Figure 1. Persistence parameter estimates: 5-year rolling window



(9) Australia



(10) Czech Republic



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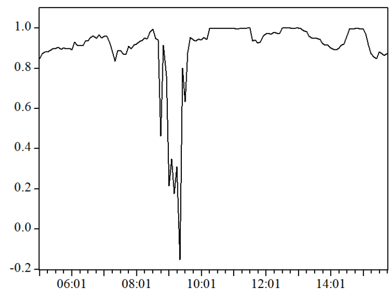
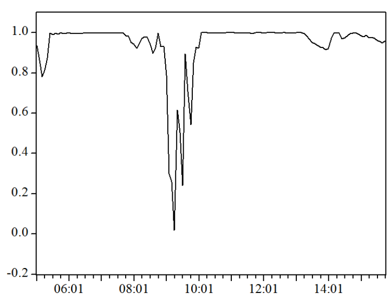
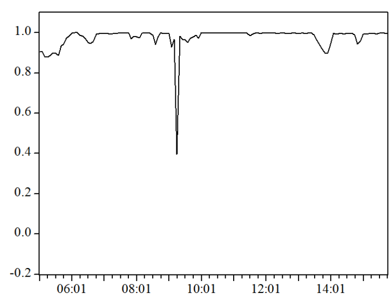


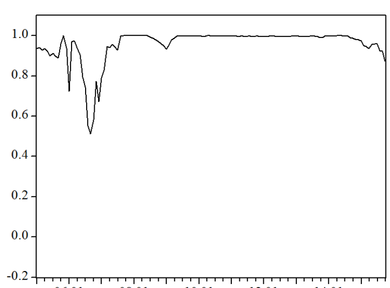
(12) Estonia



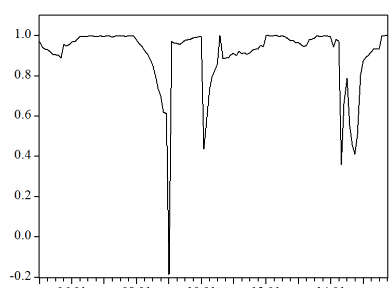
(13) Finland



(14) France

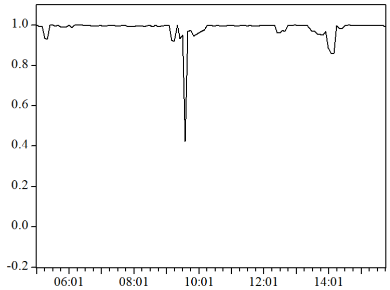


(15) Germany

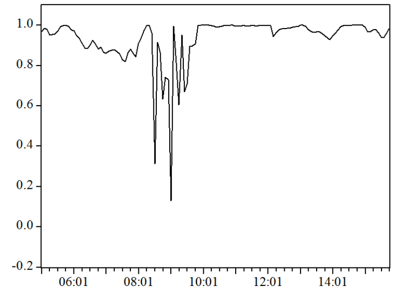


(16) Greece

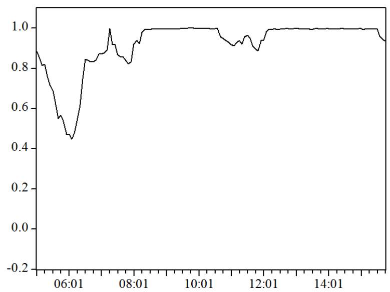
Figure 1. (*continued*)



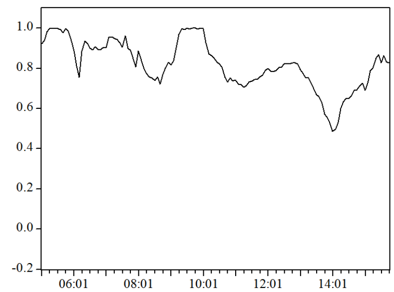
(17) Hungary



(18) Ireland



(19) Italy



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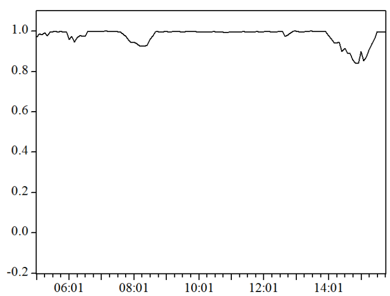
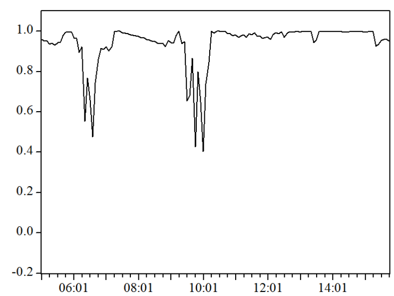


(21) Luxembourg



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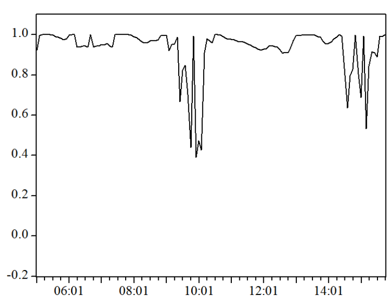
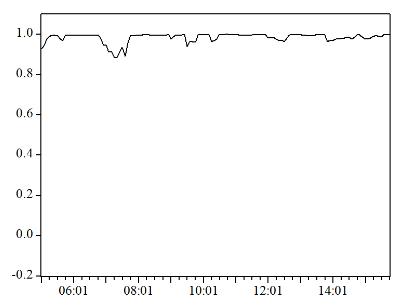
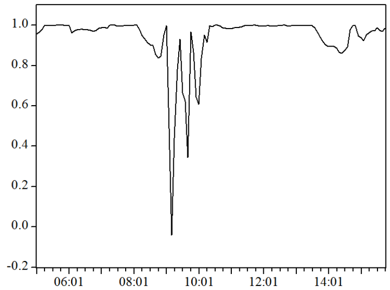


(23) Poland

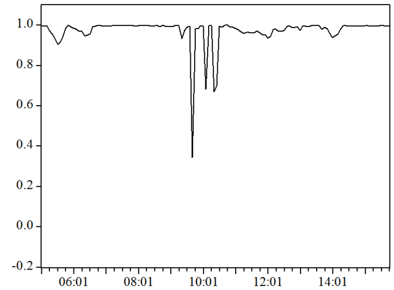


(24) Portugal

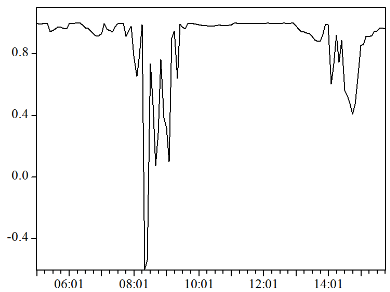
Figure 1. (*continued*)



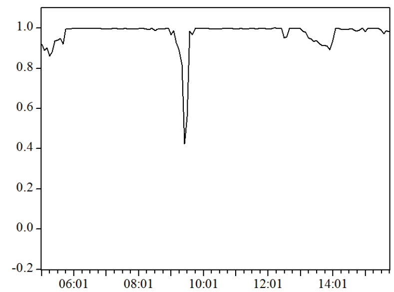
(25) Slovak Republic



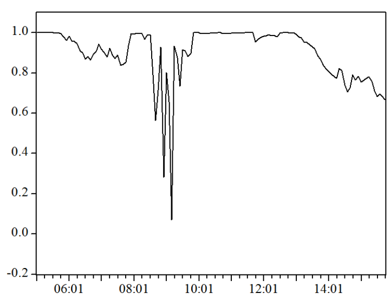
(26) Slovenia



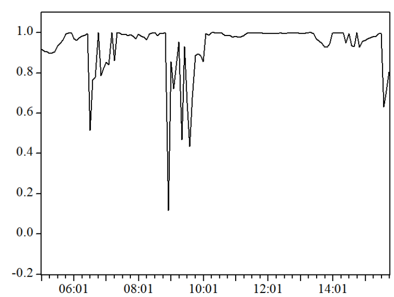
(27) Spain



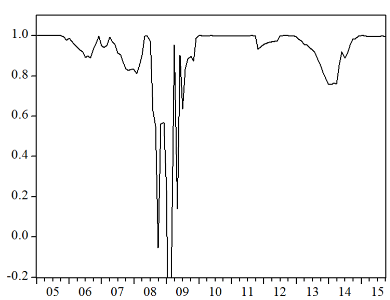
(28) United Kingdom



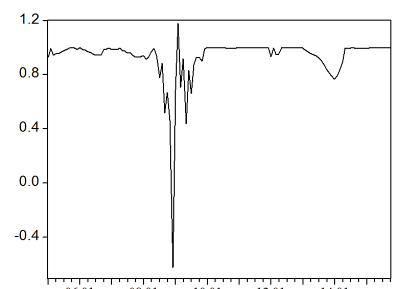
(29) United States



(30) EURO AREA



(31) G7



(32) OECD

Figure 1. (*continued*)