Department of Economics - FEA/USP

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ANDRÉ M. MARQUES

GILBERTO TADEU LIMA

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André M. Marques (00094751@ufrgs.br)

Gilberto Tadeu Lima (giltadeu@usp.br)

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JEL Codes: E24; E27.

Searching for Patterns of Unemployment Persistence in OECD Countries with Aggregated and Disaggregated Data, 2000-2014[1]

ANDRÉ M. MARQUES† and GILBERTO TADEU LIMA‡

†*Department of Economics, Federal University of Paraíba*; *Cidade Universitária, Campus I. João Pessoa, Brazil* (*e-mail: 00094751@ufrgs.br*)
‡*Department of Economics, University of São Paulo*; *Av. Prof. Luciano Gualberto, 908. São Paulo, Brazil* (*e-mail: giltadeu@usp.br*)

ABSTRACT

One major concern regarding the recent financial crisis that hit the U.S. and several other OECD countries is that it may have worsened the pattern of unemployment persistence in those countries where the rate of unemployment has remained above pre–crisis levels. In this context, we use mean bias-corrected parameter estimator and bootstrap permutation test methods with a moving window to test for changes in the pattern of unemployment persistence in 29 OECD countries before and after the recent financial crisis using monthly aggregated and quarterly disaggregated data. We estimate the most likely date of change in the trend of unemployment and use this information to compute an unbiased scalar measure of persistence to test whether the recent financial crisis has produced any significant change in the pattern of unemployment persistence. We find evidence of an increased unemployment persistence in several countries which is correlated with the recent financial crisis.

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1. Introduction

As a result of the 2007-2008 financial crisis in the United States, a wave of contractionary effects hit production and employment in several OECD countries. As recovery has been slow in several countries, there is a concern that these contractionary effects reflect changes in structural conditions that may have worsened the pattern of unemployment persistence in those countries where the unemployment rate has remained above pre–crisis levels. In fact, the two most recent editions of the OECD Employment Outlook, launched in mid-2014 and mid-2015, recognize that the persistence of high levels of unemployment appears to have been translated into a rise in structural unemployment in some countries (OECD, 2014, 2015).

In this context, we investigate patterns of unemployment persistence in the United States and other OECD countries in two different periods (before and after the 2007-2008 financial crisis). As in Cuestas et. al. (2011), we take into account the big event of the recent financial crisis as an impulse factor. In the same spirit of Banerjee et. al. (1992) and Zivot and Andrews (1992), we assume that big events and external shocks can exert its effects on economic activity in a permanent or transitory way depending on the nature and magnitude of the persistence of key macroeconomic variables, such as the unemployment rate. However, we should emphasize that it goes beyond the scope of this paper an exploration of possible fundamental causes of the observed unemployment persistence in OECD countries, a contentious issue on which there is a sizeable literature (see, e.g., Nickell, 1997; Blanchard and Wolfers, 2000; Aberg, 2000; Ball, 2009). More focusedly, our ambition is to contribute to the empirical literature by mapping out patterns of unemployment persistence in the OECD countries using both aggregated and disaggregated data for a time span before and after the 2007-2008 financial crisis.

More precisely, we employ mean bias-corrected parameter estimator and bootstrap permutation test methods with a moving window to detect possible changes in the pattern of unemployment persistence in OECD countries. We investigate patterns of unemployment persistence before and after the recent financial crisis using monthly aggregated and quarterly disaggregated unemployment data (by gender and age) for 29

OECD countries. As we estimate the most likely date of change in the trend function of unemployment, we use this information to compute an unbiased scalar measure of persistence which allows us to test (using a bootstrap permutation test) whether the recent financial crisis caused any significant change in the pattern of unemployment persistence relatively to previous periods. This unbiased scalar measure of persistence yields more accurate and useful information for stabilization policy in both the cumulative impulse response and half-life dimensions. Referring to OECD countries, some authors recognize that "[t]he 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 from the Great Recession is included. Recent events may indeed be crucial to our understanding of unemployment" (Cheng et. al., 2012, p. 429). Meanwhile, we are not aware of any other empirical investigation that computes a scalar measure of unemployment persistence taking into account explicitly the potential influence of big and rare events like the recent financial crisis on unemployment persistence in OECD countries.

Therefore, this paper contributes to the literature on unemployment persistence in OECD countries in several ways. First, in addition to using a stationary/nonstationary approach, we estimate a bootstrap mean-unbiased scalar measure of unemployment persistence for most of the OECD countries, which is absent in other studies as they are typically focused on the dichotomy between mean reverting vs. no mean reverting. Second, we use a moving window procedure of five years to make inferences about the pattern of changes in our measure of unemployment persistence over the years, for younger and older workers by gender in each individual country. Third, by considering the recent financial crisis as a big event, along with our unbiased scalar measure of unemployment persistence and the bootstrap permutation test, we find evidence that such crisis has produced a statistically significant increase in unemployment persistence in OECD countries.

We follow Pivetta and Reis (2007) and Kim (2003) in adopting a univariate time series approach to measure the magnitude of persistence in macroeconomic variables, but take several steps further. First, we apply the Perron and Rodríguez (2003) procedure to make inferences about the very 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 moving window of five years to follow the path of the unemployment persistence over time. This approach offers valuable information about smooth or abrupt changes in the coefficient which measures persistence over time, in the presence of big events such as the recent financial crisis. In fact, the moving window procedure offers a dynamic view of unemployment persistence in the aggregated and disaggregated data by gender and age for OECD countries. Some interesting patterns in the gender and age categories emerged in each country, expectedly showing their country-specific characteristics. Third, our main contribution lies in the making of inferences about the behavior of the scalar measure of unemployment persistence for full and partitioned samples before and after the recent crisis using both monthly aggregated data and quarterly disaggregated data by gender and age in each country. In the specification of the model, the moving window procedure shows that there is not much change in unemployment persistence when only quarterly data are used. Fourth, as we measure persistence with an unbiased scalar before and after the recent crisis, we follow Efron and Tibshirani (1993) and use a bootstrap permutation test to investigate whether such crisis has raised the magnitude of unemployment persistence in OECD countries. As it turns out, we can reject the hypothesis of equal persistence before and after the crisis at 10% of probability.

The remainder of the paper is organized as follows. The next section briefly places our contribution in the context of the related empirical literature on unemployment persistence. Section 3 describes the methods and data used in the estimations. Section 4 presents and discusses the main empirical results, whereas Section 5 offers final remarks.

2. Some related empirical literature

Persistent unemployment has been a dominant feature in OECD countries in the last decades. In fact, average unemployment rates for representative samples of OECD countries have increased since the early

1970s, although with considerable cross-country dispersion. For individual countries, meanwhile, patterns of unemployment usually differ for different groups of workers stratified by age and gender.

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. The authors adopt a variety of approaches to study the nature of unemployment persistence, including fractional integration analysis, unit root with and without structural breaks or with non-linear adjustments, and impulse response functions. The authors find that the unemployment rate appears to be a non-mean reverting process in most of the countries when no structural breaks are considered. When structural breaks are considered, it is only in Lithuania that the unemployment rate appears to be a mean reverting process. Meanwhile, when a fractional approach is used, the results of the preferred model (ARFIMA (1, *d*, 0)) show that in all cases unemployment persistence is very high, and similar to the results found in the present paper (as will be shown later). When the impulse response functions are considered, all these conclusions are confirmed in all cases, with the half-life measure extending for a period of eight to nine years on average.

Jiménez-Rodríguez and Russo (2012) follow a vector autoregression methodology to investigate the behavior of unemployment persistence (or responsiveness to shocks) in five countries which pursued a partial labor market reform program during the 1990s, namely, Italy, Germany France, Spain and United Kingdom. Results from the impulse response function are obtained for both before and after the reforms by splitting the sample data (a quarterly dataset on employment and output for 1980-2008). The authors find that such 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 recent crisis. 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. As their methodology incorporates recent developments in panel data analysis, the authors consider the presence of cross-section dependence among the U.S. states and obtain estimates of the half-life and moving window procedure which are absent in other studies. When more recent data are included and cross-section dependence is considered, they find strong evidence of hysteresis, with the persistence of the common component of the data tending to be very high even when there is strong evidence of stationarity, as documented by a half-life of six to fourteen years. These results lead the authors to question the practical usefulness of the natural rate hypothesis which postulates stationarity, since the half-life magnitude is quite long when compared to the typical duration of economic recessions. As shown later, we obtain similar results for the pre-crisis period, as we obtain a half-life estimate of about six years for the U.S. economy using monthly data. For the post-crisis period, however, our data and methodology generate different results for the U.S. The authors' results from the moving window procedure are also similar to ours, as they find a sudden increase in unemployment persistence in the aftermath of the recent financial crisis. Fosten and Ghoshray (2011) also adopt a mean reverting vs. non mean reverting approach, and offer a comprehensive review of the unit root tests applied to the unemployment rate. However, their results are not directly comparable to ours because they use lower frequency, annual data on unemployment, which tends to smooth the autoregressive coefficient.

3. Methodology

The data used in this paper was obtained from OECD Statistics in harmonized unemployment series. Four steps were adopted in this study. Initially, we seek evidence of stationary behavior with a more powerful GLS detrending approach to unit root test allowing occasional changes in the trend function of the series. Then, when the series exhibit stationary behavior, we split the sample into two parts to compare the unemployment persistence before and after the recent financial crisis. The procedure of moving window is the same used by Pivetta and Reis (2007), but we depart from those authors by applying the more efficient bootstrap mean unbiased estimator for the autoregressive coefficient in the AR model introduced by Kim (2003). Finally, we calculate the achieved significance level (ASL) by performing a bootstrap permutation test which enables us to reject the null hypothesis of equal unemployment persistence before and after the recent financial crisis at 10% of probability.

3.1 The dataset

The dataset used in this paper covers the period from 2000:01 to 2014:10 in most cases, yet the final period is not the same for all countries. The selection of this dataset takes into account the availability of a monthly series until January, 2014. The general idea is to select a sample which has been influenced by the recent financial crisis in the U.S. economy, the big event of the 2000s. Another important feature of this paper is the search for a pattern of persistence in disaggregated data by gender and age. Nevertheless, since for some specific countries and periods there were no observations available, Table 1 below presents in detail the country and also the period for which data were available, with the respective sample size (*n*) for the estimations.

TABLE 1

Sample of countries – several periods – in all cases the initial period is the same (*January*, *2000*).

Country/End period	*n*	Country/End Period	*n*	Country/End Period	*n*
1. Australia/2014:11	179	11. Germany/2014:11	179	21. Norway/2014:10	178
2. Austria/2014:11	179	12. Greece/2014:09	177	22. Poland/2014:11	179
3. Belgium/2014:11	179	13. Hungary/2014:10	178	23. Portugal/2014:11	179
4. Canada/2014:12	180	14. Ireland/2014:11	179	24. Slovak Republic/2014:11	179
5. Chile/2014:10	178	15. Italy/2014:11	179	25. Slovenia/2014:11	179
6. Czech Republic/2014:11	179	16. Japan/2014:11	179	26. Spain/2014:11	179
7. Denmark/2014:11	179	17. South Korea/2014:11	179	27. Sweden/2014:11	179
8. Estonia/2014:10	178	18. Luxembourg/2014:11	179	28. United Kingdom/2014:10	177
9. Finland/2014:11	179	19. Mexico/2014:11	179	29. United States/2014:12	180
10. France/2014:11	179	20. Netherlands/2014:11	179	------------	----

Source: Authors' elaboration based on data extracted from OECD Statistics.

For monthly data, the periods are the same as those specified in Table 1, and for quarterly base data information is available from 2000 Q1 to 2014 Q3, which amounts to about sixty observations for younger and older workers for both male and females.

3.3 Econometric models and the measures of persistence

3.3.1 Testing for infinite persistence with occasional change

A large proportion of the literature follow the original work of Nelson and Plosser (1982) on theoretical and empirical issues on persistence in macro time series. Those authors conclude that macro time series of the U.S. economy were best characterized as an infinite persistence process, arguing that current shocks have permanent effects for most of the fourteen macro series analyzed. Perron (1989) argues that those results change if a more flexible trend function is adopted, as there is greater evidence against the null of a unit root. Zivot and Andrews (1992) and Banerjee et. al. (1992) (BLS) criticized the Perron (1989) approach, arguing that it was necessary to adopt a data-dependent algorithm which estimates the time of changes that maximizes the *t*-statistic of the test, independently from any prior information of the researcher. They analyze these series by adopting a flexible trend function with one occasional change in an unknown time and find less evidence against the null hypothesis of infinite persistence, thus partially supporting the previous conclusions reached by Nelson and Plosser (1982).

Perron and Rodríguez (2003) recognize that ADF and PP tests suffer from low power to detect the persistence in time series in the context of trend shift or big events in the economy. A more efficient approach was introduced in the literature of unit root tests, the so called GLS detrending approach (Elliot et. al., 1996). By allowing occasional change in the trend function in an unknown time, based on the local to unity GLS detrending approach, in this paper we test the null hypothesis of infinite persistence with two specific models: model I allows for a break in the slope of the trend function and model II allows for a break in both the intercept and the slope. We choose this methodology by considering the advantages of the class of M-tests because these tests have much smaller size distortion than other classes of unit root tests when the errors have strong negative serial correlation. Moreover, the use of GLS detrending when constructing the M-tests allows higher gains in power similar to the DF^{GLS} test introduced by Elliot et. al. (1996). The basic model in this approach is given by:

$$y_t = d_t + u_t, t = 0, ..., T \quad (1)$$

And:

$$u_t = \alpha u_{t-1} + \varepsilon_t, \quad (2)$$

where ε_t is an unobserved stationary process with zero mean. In (1), $d_t = \psi' z_t$, where z_t is a set of deterministic components. Therefore, for a time series of unemployment y_t, with deterministic components z_t, the transformed data $y_t^{\bar{\alpha}}$ and $z_t^{\bar{\alpha}}$ are defined by:

$$y_t^{\bar{\alpha}} = [y_0, (1-\bar{\alpha}L)y_t] \quad (3)$$

And:

$$z_t^{\bar{\alpha}} = [z_0, (1-\bar{\alpha}L)z_t],\ t = 0, ..., T. \quad (4)$$

Then, we get $\hat{\psi}$ to be the estimate that minimizes,

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

For a structural change in the slope, the set of deterministic components, z_t in (.) is given by:

$$z_t^{\bar{\alpha}} = [1, t, I(t > T_b)(t - T_b)] \quad (6)$$

where $I(\cdot)$ is the indicator function and T_b is the time of change; it is assumed that $T_b = T\delta$ for some $\delta \in (0,1)$. In this model, $\psi'(\delta) = (\hat{\mu}_1, \hat{\beta}_1, \hat{\beta}_2)'$ is the vector which minimizes (5). For a structural change in the intercept and slope, we have:

$$z_t^{\bar{\alpha}} = [1, I(t > T_b), t, I(t > T_b)(t - T_b)] \quad (7)$$

The asymptotic distribution of the test explores the feature that a series converges with different rates of normalization under the null and the alternative hypothesis, whose statistic of the test is defined by eq.(8) in Perron and Rodríguez (2003, p. 4). In this case the interest lies in the ADF test, denoted by $\text{ADF}^{\text{GLS}}(\delta)$, which results in the t-statistic to test if $b_0 = 0$ in the following regression:

$$\Delta\tilde{y}_t = b_0\tilde{y}_{t-1} + \sum_{j=1}^{k} b_j \Delta\tilde{y}_{t-j} + e_{tk}, \qquad (8)$$

where $\tilde{y}_t = y_t - \hat{\psi}' z_t$.

3.3.2 A scalar measure of persistence: bootstrap mean bias-corrected estimator

In this approach, persistence has the meaning of long-run effect of a shock to the level of unemployment – given a shock that raises unemployment today by 1%, by how much do we expect it to be higher at some future date? Our approach to the issue follows similar studies by Taylor (2000) and Pivetta and Reis (2007). However, unlike those authors, to express the magnitude of persistence we use the bootstrap mean bias-corrected estimator introduced by Kim (2003). It showed better performance in simulations with respect to forecast accuracy against existing alternatives in the literature, when AR root is close or equal to one especially in small samples. In these simulations with small samples, the performance of the bootstrap mean bias-corrected estimator was compared with the median unbiased estimator, introduced by Andrews and Chen (1994) and the bias corrected estimator introduced by Roy and Fuller (2001). There are two main differences between using these approaches and adopting the algorithm proposed by Kim (2003). First, the bootstrap mean-bias corrected estimator is made explicitly conditional on the last p observations of the series. Second, this estimator corrects for biases in all parameters in the model simultaneously and this is different from Andrews and Chen (1994) and Roy and Fuller (2001). Thus, following Andrews (1993) and Kim (2003), our baseline model was specified as an *augmented Dickey-Fuller regression*, given by:

$$Y_t = \mu + \beta t + \alpha Y_{t-1} + \psi_1 \Delta Y_{t-1} + \ldots + \psi_{p-1} \Delta Y_{t-p+1} + U_t, \quad (9)$$

for $t = 1,\ldots,T$ where $\{Y_t : t = -p+1,\ldots,T\}$ is the series of unemployment, and $U_t \sim iid\left(0,\sigma^2\right)$ for $t = 1,\ldots,T$. Thus, we have the downward biased vector of estimated coefficients $\hat{\gamma} = \left(\hat{\mu},\hat{\beta},\hat{\alpha}\right)$ which must be corrected by the procedure decribed below. The central feature of this specification is that if $\alpha \in (-1,1)$ the cumulative effect of a shock in unemployment rate is given by $\dfrac{1}{1-\alpha}$, so that this magnitude gives the total cumulative effect of a unit shock on the entire future of the time series. Then, for a larger α we have higher persistence of unemployment. As a consequence, by seeking the best AR(*p*) model for each country we can estimate this measure for unemployment persistence before (b) and after (a) the recent financial crisis to test whether this change in parameters is significant or not by adopting a bootstrap permutation test introduced by Efron and Tibshirani (1993). Moreover, we perform the same estimation for four classes or groups of workers with disaggregated quarterly data by gender and age. In summary, to correct simultaneously for biases in all coefficients of the model, we adopt the following steps: (1) the Y_t^* random sample is generated; (2) we regress Y_t^* against $\left[1, t, Y_{t-1}^*, Y_{t-p}^*\right]$; (3) as a result, we obtain $\gamma^* = \left(\mu^*, \beta^*, \alpha^*\right)$; (4) the bias is calculated as $bias\left(\hat{\gamma}\right) = \overline{\gamma}^* - \hat{\gamma}$, in which $\overline{\gamma}^*$ is the sample average; and finally (5) we obtain the mean bias corrected estimates as $\hat{\gamma}^c = \hat{\gamma} - bias\left(\hat{\gamma}\right)$. We repeat this procedure 1,000 times.

There is a wide consensus about the downward bias of the AR(1) coefficient of the Least Square Estimator when the specification of the model includes drift and trend (see Andrews, 1993; Kim, 2003). However, according to Andrews (1993), this bias seems to be absent when we specify a pure AR(1) without either drift or trend. To partly control for it, in the moving window procedure, we specify the AR(1) model only with drift.

This latter specification explicitly allows for changes in unemployment persistence when the estimates are obtained through rolling regressions for a specified lag *p*, which is our first *visual measure* of persistence. This procedure offers a good idea of the *direction of changes* in the path of persistence over time. However, our main measure of unemployment persistence is the sum of coefficients, the α parameter in equation (9), which is obtained by the same procedure described in Kim (2003). This is important because it is linked to the cumulative impulse response function mentioned above, and also because this measure can provide the length of time (in this study, in months for aggregated and quarters for disaggregated data) until the impulse response of a unit shock is equal to half of its original magnitude: the *half-life of a unit shock* is defined by $HL = \left| \frac{\log(1/2)}{\log(\alpha)} \right|$. This measure characterizes the most probable duration of unemployment in all countries of the sample.

It should be noted that the moving window procedure adopts, as a measure of persistence, only the first autoregressive coefficient of our model (9) for a fixed value of p=1. We expect to find values between 0 and 1 for low persistence unemployment series and values near to or higher than one in other cases; with the moving window procedure we visualize the *changes of coefficient* over the years and this information can be carried out until our last specification of the AR(p) model for disaggregated data. It shows the direction of changes and the contrast before and after the recent financial crisis and can inform the eventual changes in parameter of the model. Since there is a recognized downward bias in the autoregressive coefficient, in the moving window procedure we do not intend to infer the precise magnitude of persistence in view of its bias in high persistent series or near unit root. At first, we use this measure to point out the *direction* of movement of persistence over the years and identify long swings (abrupt or smooth) in the α coefficient.

Therefore, as the financial crisis hit the U.S. and several OECD countries around 2007-2008, unemployment persistence is expected to have gone through changes that can be easily visualized with this procedure. The estimation starts with the 2000:01-2004:12 period, with increases of one month in each estimate of the sample and dropping the first observation (it fixes n = 60 in each estimation). Thus, the second coefficient is estimated for the 2000:02-2005:01 period, and so on, generating about 122 coefficients at the end, which are then ordered in time and showed in the Appendix. The same procedure is applied to quarterly disaggregated data by age and gender for each country.

3.3.3 Permutation test for difference in unemployment persistence before and after the financial crisis

Finally, when we obtain the unbiased scalar measure of unemployment persistence, we employ a bootstrap permutation test to verify whether the measures of unemployment persistence in OECD countries have changed significantly from before to after the recent financial crisis. This test is discussed in Efron and

Tibshirani (1993, Chap. 15). The idea is simple, and, more importantly, is free of mathematical or behavioral assumptions, since it works only with empirical distributions of the samples.

In our case, we have two independent random samples, $\mathbf{x} = (x_1, ..., x_m)$ and $\mathbf{z} = (z_1, ..., z_n)$, which are drawn from possibly different distributions *F* and *G*, and we want to test the null hypothesis of $\mathbf{H_0}$: $F = G$. This means that *F* and *G* assign equal probabilities to all sets: $\text{Prob}_F\ \{A\} = \text{Prob}_G\ \{A\}$ for A any subset of the common sample space of the $x_s^{'}$ and $z_s^{'}$. Then, if H_0 is true, there is no difference between the probabilistic behavior of a random variable *z* or *x*. For our two sample problem here, the difference between the means is:

$$\hat{\theta} = \bar{x} - \bar{z} \qquad (10)$$

We are assuming that if the null hypothesis H0 is not true, we expect to observe larger values of $\hat{\theta}$, where $\bar{x} - \bar{z}$ is the difference in persistence of unemployment after and before the recent financial crisis. If the presumed influence is large, we expect this $\hat{\theta}$ to be large as well. Given the observed value of $\hat{\theta}$, the achieved significance value (ASL) of the test is defined as:

$$\text{ASL} = \text{Prob}_{H0}\left\{\hat{\theta}^* \geq \hat{\theta}\right\} \qquad (11)$$

Hence, the smaller the value of ASL, the stronger the evidence against H_0 is. Since ASL can be interpreted as a degree of credibility of H_0, low credibility suggests that we should reject it. Note that $\hat{\theta}$ is a fixed value and $\hat{\theta}^*$ has the null hypothesis distribution and will be generated according to it. The logic is that if the null hypothesis is correct, then any of the persistence measures for any country could have come equally well from either one of the samples we are testing. The test combines all the *m* plus *n* observations, where $N = m + n$ in both groups of countries together, then take a random sample of size *m* of the first group without replacement; the remaining *n* observations constitute the second group. We compute the difference between the mean for each group and then repeat this procedure 1,000 times. Formally, there are $\binom{N}{n}$ permutation replications of $\hat{\theta}^*$ and the distribution that attaches probability $\frac{1}{\binom{N}{n}}$ to each of these is called permutation distribution of $\hat{\theta}$ or $\hat{\theta}^*$. The ASL permutation is defined as:

$$\text{ASL}_{\text{perm}} = \text{Prob}_{\text{perm}}\left\{\hat{\theta}^* \geq \hat{\theta}\right\} = \frac{\#\left\{\hat{\theta}^* \geq \hat{\theta}\right\}}{\binom{N}{m}} \qquad (12)$$

where # indicates the number of times of the realized values inside the keys.

4. Results and discussion

In this section, the results are reported for the four stages of the empirical investigation, namely: a) tests for stationarity/nonstationarity, by showing the most likely time break in the path of unemployment; b) estimation of unemployment persistence for the full sample and partitioned ones; c) analysis of the temporal evolution of the measure of unemployment persistence for a 5-year moving window, and d) presentation of unbiased estimates and test for changes in persistence with a bootstrap permutation test.

4.1 Evidence on stationarity and partition of the sample

As emphasized in Andrews (1993), the traditional ADF and PP tests in general have low power against the alternatives of high persistence. Moreover, these tests do not take into account possible structural changes in the trend function of the series. As argued in Perron (1989) and Zivot and Andrews (1992), by ignoring possible trend breaks, the traditional unit root tests may lead to incorrect conclusions. In this paper we assume that this possible break date, estimated through the Perron and Rodríguez procedure, may serve as a valuable indicator of change in the degree of persistence of the series. We follow a strategy similar to that adopted by Okimoto and Shimotsu (2010) in partitioning the sample to verify whether there are significant differences between two situations: before and after the recent financial crisis.

Yet, unlike Okimoto and Shimotsu (2010), we estimate the most probable date of structural change in the trend function of unemployment by applying the Perron-Rodríguez (2003) unit root test, in which the date of the break is selected by maximizing the absolute value of the *t*-statistic on the dummy variable in both Models I and II, as explained previously. With this procedure, we expect that changes in our measure of persistence will be more correlated with the global financial crisis that hit most of our sample of countries. Recall that Model I allows for structural change in the slope only and Model II allows for structural change in the intercept and slope. Tables 2-6 below present the results for the $MADF^{GLS}$ tests choosing the date for the largest absolute value of the *t*-statistic for the trend break dummy.

In Table 2, we report the results for the aggregated unemployment rate, for our sample of 29 countries. Since the United States is the main seed and motivation for this empirical study, we apply both the Perron and Rodríguez test and the Banerjee et. al. (1992) (BLS) unit root test, which yield different results. To reach more confidence in the conclusions we also apply the Zivot and Andrews (1992) unit root test for the U.S. unemployment rate by allowing for a change in both the intercept and the slope. This procedure (by fixing the lag length of 8, following Zivot and Andrews (1992)) yield the same date as the BLS test in 2008:04 and enable us to reject the null hypothesis of infinite persistence at 5% and 10% of probability.[2] The BLS and Zivot and Andrews (1992) results of unit root tests for the U.S. economy seem more correlated with the financial crisis emanated from that country. The main conclusion to draw from the results presented in Table 2 is that we can reject the null hypothesis of infinite unemployment persistence in almost all cases, except in eight countries: Austria, Belgium, Canada, Chile, Japan, Mexico, Norway and Sweden. When analyzing the aggregated unemployment data, it seems that unemployment is a mean reverting process in the majority of countries (72% of the sample). The question that remains is*: how long does it take for this to occur*?

After these aggregated initial results, as a first look on the path of persistence over the years, we present the paths of the autoregressive coefficient estimated by moving window procedure in each country in the Appendix. As noted previously, this procedure does not provide the precise magnitude of persistence, but

[2] Critical values for the Zivot and Andrews unit root test are: 0.01= -5.57; 0.05= -5.08; 0.1= -4.82 and t-statistic calculated was -5.1904.

it yields valuable information about its evolution over time and *direction of changes*. In the end, we proceed to look for the best AR(*p*) model for each weak stationary unemployment rate based on the results of Table 2 above for the remaining nineteen countries of the sample and compute their magnitude of persistence, since the AR(*p*) model yields less biased estimates of the autoregressive coefficients AR(1).

Since there is not enough data for two countries with partitioned samples [NA], we exclude Hungary and Netherlands from our potential stationary unemployment rates. From the best AR(*p*) model for each country and the sum of autoregressive coefficients, we can test for changes in persistence before and after the U.S. financial crisis in nineteen countries where the path of unemployment can be reasonably considered a stationary series with a change in the trend function in a probable date.

By considering the results reported in Table 2 above and the level of unemployment in the last year of the sample, interesting contrasts can be noted. In the countries where the level of aggregated unemployment is currently relatively higher (Greece, Ireland, Portugal and Spain), the persistence of unemployment may be considered finite and lower than in those cases where we observe an unemployment rate which is currently relatively lower and may be considered with infinite persistence (Canada, Austria, Belgium and Norway). Overall, using unemployment aggregated data for the 29 countries of the sample, we can reject the null of infinite persistence for 21 (72%) of them.

TABLE 2

Searching for structural break in the path of unemployment – results for aggregated data.

	$\mathbf{MADF^{GLS}}: H_0 : Q_t^i \approx I(1),\ i = 1, \ldots, 29.$				
Countries	N	**Model I** (*t-statistic*/*break date*)	$\hat{k}$	**Model II** (*t-statistic*/*break date*)	$\hat{k}$
1. Australia	179	[1:98] 9.1842 / 2008:02[a]	1	9.2120 / 2008:02[a]	3
2. Austria	179	2.6693 / 2004:09	2	2.8987 / 2005:08	2
3. Belgium	179	1.7732 / 2012:03	3	2.1934 / 2004:04	3
4. Canada	180	1.9836 / 2012:01	0	2.3404 / 2013:03	0
5. Chile	178	3.6947 / 2002:09	6	3.9623 / 2002:11	6
6. Czech Republic	179	[1:99] 6.3498 / 2008:03[a]	4	6.5937 / 2008:03[a]	4
7. Denmark	179	[1:88] 5.7922 / 2007:04[a]	4	5.8004 / 2007:04[a]	4
8. Estonia	178	[1:74] 5.5183 / 2006:02[a]	4	5.5289 / 2006:05[a]	4
9. Finland	179	[1:99] 15.0836 / 2008:03[a]	2	15.0221 / 2007:11[a]	2
10. France	179	[1:98] 13.3014 / 2008:02[a]	1	13.2706 / 2007:11[a]	1
11. Germany	179	[1:65] 23.6983 / 2005:05[a]	2	23.5074 / 2005:09[a]	3
12. Greece	177	[1:106] 28.5431 / 2008:10[a]	2	28.2971 / 2008:05[a]	2
13. Hungary	178	[NA] 9.4927 / 2013:02[a]	1	9.0081/2013:01[a]	3
14. Ireland	179	[1:62] 6.9763 / 2005:02[a]	6	6.9417 / 2005:04[a]	6
15. Italy	179	[1:95] 21.6739 / 2007:11[a]	1	21.5673 / 2007:08[a]	1
16. Japan	179	3.0095 / 2012:04	3	3.2072 / 2011:10	3
17. South Korea	179	[**1:30**] 4.9941 / 2002:06[a]	4	5.2326 / 2003:01[a]	4
18. Luxembourg	179	[1:56] 5.8868 / 2004:08[a]	3	5.9976 / 2005:03[a]	3
19. Mexico	179	2.8600 / 2010:10	2	3.1798 / 2012:02	2
20. Netherlands	179	[NA] 7.0382 / 2011:06[a]	3	6.8143 / 2011:03[a]	3
21. Norway	178	2.9585 / 2003:08	4	3.3112 / 2004:01	4
22. Poland	179	[**1:30**] 9.0321 / 2002:06[a]	4	8.6784 / 2002:10[a]	4
23. Portugal	179	[1:100] 4.3727 / 2008:04[c]	2	4.6527 / 2012:12[b]	2
24. Slovak Republic	179	[1:105] 8.4950 / 2008:09[a]	1	8.5597 / 2008:07[a]	1
25. Slovenia	179	[1:105] 13.3464 / 2008:09[a]	4	13.2647 / 2008:03[a]	5
26. Spain	179	[1:82] 21.9767 / 2006:10[a]	2	21.9184 / 2006:11[a]	2
27. Sweden	179	0.9282 / 2010:04	3	1.5647 / 2012:05	3
28. United Kingdom	177	[1:49] 6.4882 / 2004:01[a]	4	6.4626 / 2004:10[a]	4
29. United States	180	BLS[**2008:04**] 8.0958 / 2011:09[a]	5	8.3343 / 2010:11[a]	5

Note: Critical values for $\mathbf{MADF^{GLS}}$: **Model I** = (1%, 5%, 10%) **-4.87**; **-4.33**; **-3.99** and **Model II** = (1%, 5%, 10%) **-4.86**; **-4.48**; **-4.02**. All critical values were extracted from Perron and Rodríguez (2003, p. 11, Tab. 1b) for N=100. (**a**) Significant at 1% probability; (**b**) significant at 5% probability; (**c**) significant at 10% probability. The number of additional lags (k) was obtained by the sequential *t-test* at 10% of significance because in the case of applying $\mathrm{MADF^{GLS}}$ for all data-dependent methods, according to the authors "the power is high for all methods to choose k" (Perron and Rodríguez, 2003, p. 13). See Perron and Rodríguez (2003, p. 12-13) for details of the procedure, results of simulations and comparisons among the three different criteria.
Source: Authors' calculations from the dataset provided by OECD Statistics.

We now present the results of the Perron and Rodríguez (2003) test for quarterly disaggregated data by gender and age for the same period of time (2000:1 to 2014:4). In this case, in each country, we consider four categories of workers: (1) male, 15-24 years old; (2) male, 25-54 years old; (3) female, 15-24 years old; and (4) female, 25-54 years old. Admittedly, this categorization does not include all the people who are willing to work in the economy, but we are using the available data in OECD statistics and this does not exclude other possibilities.

Table 3 presents the results for the first group of workers (male, 15-24 years old), and the following main conclusions emerge. First, in many countries of the sample, the estimate break date seems to correlate with the recent financial crisis in the U.S. However, when we compare these results with those obtained using aggregated data, there is less evidence against the null hypothesis of infinite persistence. Second, by using Model I, there are only eleven countries (44% of the sample) where we may consider the rate of unemployment for young male as stationary. This result is in stark contrast with the equivalent one obtained using aggregated data, in which we can reject the same null hypothesis in 72% of the cases.

[TABLE 3 AROUND HERE]

Table 4 presents results for the second group of workers: 25-54 year old male, who are expected to be more experienced and skilled than young workers.

[TABLE 4 AROUND HERE]

For this category of workers we have a greater proportion of cases where we can reject the null hypothesis of infinite persistence when compared to the first group of less experienced workers. In fact, in 52% of the cases the unemployment rate can be considered as stationary.

Table 5 presents the results of the Perron-Rodríguez test for the third group of workers: female, 15-24 years old.

[TABLE 5 AROUND HERE]

The most noticeable result which emerges from this disaggregated quarterly data analysis seems to be the one presented in Table 6 below: in 68% of the countries there is strong evidence against the null hypothesis of infinite persistence for female, 25-54 years old workers.

[TABLE 6 AROUND HERE]

4.2 Moving window procedure: discussion

Next, we consider the results of our procedure of moving window for aggregated and disaggregated data. The corresponding figures presented in the Appendix show the changes in persistence over the years. F**igure 1a** shows the path of unemployment persistence for four countries where the unemployment rate can be considered as nonstationary (Austria, Belgium, Canada and Chile). In spite of having different behaviors, note that they show great dispersion of unemployment persistence around 2008, when Canada seems to be the most affected country in this group. This seems reasonable, since it has a stronger relationship with the U.S. economy. Overall, there are no general trends in higher or lower persistence over the years in these countries. In **Figure 1b** (Japan, Mexico, Norway and Sweden) we note a temporary trend of lower persistence since 2006, and, after the 2008 crisis in all these four countries there is a jump in the coefficient. The most affected country is Mexico, and likely for the same reason as Canada. We observe, in all these four countries, a trend of reduction in persistence nearly 2014, except in Japan, whose persistence has grown since 2008 and has been steady over the years.

In **Figure 1c** (Australia, Czech Republic, Denmark and Estonia) we note a jump in the persistence of unemployment in the Czech Republic and Estonia around 2008 and 2010. After this latter year, there was an apparent reduction on persistence in all four countries, until a new wave of growing persistence after 2012. In spite of it, Czech Republic seems to be the only country in this group where there is a high and stable persistence of unemployment over the years.

In **Figure 1d** (Finland, France, Germany and Greece) we observe an apparent dispersion in the persistence of unemployment in all four countries. However, the 2008 financial crisis seems to produce synchronization in those labour markets, because the dispersion after the crisis was substantially reduced and turned into a common path. However, the only country where we observe higher and more stable persistence in relation to the others is Greece. In **Figure 1e** (Hungary, Ireland, Italy and South Korea) it seems that the main affected countries are Ireland and South Korea. Hungary and Italy, even before the recent financial crisis, already pertained to sub-group of countries with high and stable unemployment persistence. However, in spite of a lower initial level of persistence, the South Korean labor market has reduced its unemployment persistence much more over the years. In contrast, in countries like Ireland, Italy and Hungary, we observe the same pattern: a high and stable persistence over the years.

In **Figure 1f** (Luxembourg, Netherlands, Poland and Portugal) we observe that Poland was the first of these four countries to experience a considerable increase in the measure of unemployment persistence (near 2006). However, in all other countries, the 2007-2008 financial crisis seems to have produced a growing and common trend in unemployment persistence until stabilization in high levels over the years, near to unity. In F**igure 1g** (Slovak Republic, Slovenia, Spain, United Kingdom, United States) we note a considerable dispersion before the 2007-2008 financial crisis, and, after the crisis an apparent synchronization process. The most pronounced and synchronized jump occurred with Spain, U.S. and U.K.

Let us now turn to some interesting results obtained using quarterly disaggregated data, which are also pictured in the Appendix. First, in Australia, the group of experienced female workers (25-54 years old) has the lowest persistence in relation to other groups in the same country, and this behavior seems to have become more pronounced since 2011. In contrast, Austria presents the same behavior, even for young female (15-24 years old). This downward path of persistence was observed in that economy since 2011 as well. In Belgium, we observe the same behavior for younger female workers (15-24 years old), who present the lowest persistence in unemployment and there was also a noticeable downward trend and stabilization since 2011. In this latter country, a contrasting behavior appears for female workers (25-54 years old), for whom we note a growing trend since 2011.

Second, interestingly, there is little intra-group (by age and gender) difference in unemployment persistence in the following countries: Canada, Czech Republic, Hungary, Ireland, Israel, Italy, Netherlands, New Zealand, Poland, Portugal, Slovak Republic, Spain, United Kingdom and United States. However, these plots reveal an interesting phenomenon: there is great dispersion in unemployment persistence among groups before the recent crisis in Italy and Spain, but these countries experience a domestic inter-group convergence in terms of persistence after that, and to a higher and unprecedented level.

Third, in Denmark we observe a marked difference between two groups. For male workers (25-54 years old) and female workers (25-54) there is a higher and more stable persistence over the years: more experienced workers have the same characteristics. And, for younger workers of both genders persistence of unemployment seems lower and similar. This contrasting behavior is observed in Estonia as well, but in a different fashion: for younger workers (males, 15-24 years old and females, 15-24 years old) there is a low and stable persistence over the years; but, for more experienced workers of both genders there is higher and stable persistence over the years. This kind of dual behavior in terms of unemployment persistence is also observed in Finland, Norway, Estonia and Slovenia. Finally, in Japan and South Korea we find a break in persistence since 2009 for younger females (15-24 years old) in the former country, and for male workers (15-24 years old) in the latter country.

4.3 Bootstrap mean bias-corrected estimator: results

After mapping out the nature of unemployment persistence using monthly aggregated and quarterly disaggregated data by age and gender, we now proceed to estimate the *magnitude* of this key property of the unemployment series. Table 7 presents the results for the scalar measure of persistence introduced by Kim (2003) in a AR(*p*) model. It consists in the bootstrap mean bias-corrected estimates for the α parameter in our baseline model (9).

Two scalar measures were computed to express the persistence of unemployment in the countries whose series could be considered stationary according to the Perron and Rodríguez approach: the α coefficient (sum of autoregressive coefficients) and the popular *half-life of unit shock* (HL). For $\alpha \geq 0$, the HL gives the length of time until the impulse response of a unit shock will be equal to half of its original magnitude. From the results presented in Table 7 we find that there is a substantial change in the persistence of unemployment in OECD countries by comparing our scalar measure before and after the recent financial crisis: in *ten countries* we observe a *higher persistence* for unemployment *after the recent financial crisis*. In *five countries* we note a *lower persistence* (Denmark, Finland, France, Slovak Republic and United States). The remaining *four countries* maintained the *same level* of persistence before and after the financial crisis. Moreover, noote that the magnitude of the corresponding change is relatively high in some countries. In fact, for Australia, Estonia, Ireland, Italy, Poland and United Kingdom such change is substantial. From a pre-crisis finite persistence, measured by the half-life in months, the post-crisis persistence becomes infinite. In the cases of Czech Republic, Portugal, Slovenia and Spain, note that the recent financial crisis seems not to have altered the gray picture of infinite unemployment persistence, since we observe the same magnitude value before and after the crisis.

TABLE 7

Bootstrap mean bias-corrected estimates for unemployment in OECD countries using aggregated data.

Countries	$\hat{\alpha}^b$	$\hat{\alpha}^a$	HL(b)	HL(a)	*p*(b)	*p*(a)
1.Australia	0.93	1.00	9.6	∞	4	5
2.Czech Republic	1.00	1.00	∞	∞	5	5
3.Denmark	1.00	0.96	∞	17.0	2	5
4.Estonia	0.86	1.00	4.6	∞	3	4
5.Finland	0.98	0.97	34.3	22.8	5	3
6.France	1.00	0.94	∞	11.2	2	4
7.Germany	0.95	0.98	13.5	34.3	1	4
8.Greece	0.99	1.00	69.0	∞	1	3
9.Ireland	0.95	1.00	13.5	∞	4	2
10.Italy	0.93	1.00	9.6	∞	1	1
11.South Korea	0.71	0.89	2.0	5.9	2	5
12.Luxembourg	0.94	0.98	11.2	34.3	5	1
13.Poland	0.59	1.00	1.3	∞	3	5
14.Portugal	1.00	1.00	∞	∞	3	3
15.Slovak Republic	1.00	0.96	∞	17.0	2	3
16.Slovenia	1.00	1.00	∞	∞	4	5
17.Spain	1.00	1.00	∞	∞	2	2
18.United Kingdom	0.92	0.99	8.3	69	1	2
19.United States	0.99	0.90	69.0	6.6	4	5

Note: The maximal order *p* considered was 5 in all models and the BIC criterion was used to select the best lag length. In all cases we specify the model with drift and trend, performing 1,000 bootstrap replications. HL is measured in the same time periods of data.
Source: Authors' calculations from the dataset provided by OCDE Statistics.

Table 8 summarizes the results of mean bias-corrected estimator for the persistence parameter and the half-life in quarters, using disaggregated data (male, 15-24 years old and 25-54 years old).

[TABLE 8 AROUND HERE]

In general, the difference between the means of the two groups is small (0.02). In spite of the fact that the first group of workers present a higher number of cases with infinite persistence (six cases), there seems to be no marked difference between young and older workers as a pattern for countries in general, only at specific ones. For example, in the U.S. and Greece we observe the opposite pattern: in the former, lower age is related to lower unemployment persistence and vice versa; in Greece, lower age for male workers is related to higher unemployment persistence. In the Hungarian labor market, we note a pattern similar to the U.S. one. Table 9 summarizes the results for our main measures of unemployment persistence for female workers with 15-24 and 25-54 years old.

[TABLE 9 AROUND HERE]

Some main conclusions which can be drawn from Table 9 are the following: First, like the results presented above, in general, the difference between the means of the two groups is small (0.01). Second, there are high contrasts between older and younger workers, in a specific way: in the United Kingdom, United States and Portugal, lower age is associated with higher unemployment persistence and vice versa. In some specific countries, like Hungary, Ireland, Spain and Australia, the persistence is infinite for both groups of workers. An interesting case is observed in South Korea, where, in all cases considered, it is found the lowest levels of both average unemployment and unemployment persistence.

Table 10 summarizes our main results. These results for the p-value by permutation test are obtained by using 1,000 bootstrap replications for the difference between the mean for each group, including α before and after the financial crisis, and for another four possible hypothesis of interest.

TABLE 10

Results for the permutation test for change in unemployment persistence

Null hypothesis	**ASL**
H0.1 : $\alpha^{a} = \alpha^{b}$	0.0890
H0.2 : $\alpha^{(2)} = \alpha^{(1)}$	0.6950
H0.3 : $\alpha^{(4)} = \alpha^{(3)}$	0.7240
H0.4 : $\alpha^{(4)} = \alpha^{(2)}$	0.8350
H0.5 : $\alpha^{(3)} = \alpha^{(1)}$	0.9110

Source: Authors' calculations from the dataset provided by OECD Statistics.

Taken together, our empirical results show the occurrence of a significant change in the pattern of unemployment persistence in OECD countries over the 2000-2014 period. At 10% of probability, there is borderline evidence of a change in the persistence of unemployment in these countries which is correlated with the financial crisis in the U.S. economy. Nonetheless, there is a great deal of heterogeneity in the actual experience of individual countries.

5. Final remarks

Motivated by the widespread concern that the recent financial crisis that hit the U.S. and several OECD countries may have worsened the pattern of unemployment persistence in those countries, this paper carries out a comprehensive empirical investigation of this issue using both aggregated and disaggregated data. By using a variety of econometric methods, this paper offers robust empirical evidence on the evolution of the pattern of unemployment persistence in OECD countries from 2000 to 2014. Revealingly, therefore, we can explore the potential influence of the 2007-2008 U.S. financial crisis on the evolution of such patterns of unemployment persistence. The moving window procedure shows no structural change in the value of the coefficient of unemployment persistence which has a stable trajectory, approximately constant for the lower-frequency quarterly data. However, using higher-frequency monthly data, the same procedure reveals some important swings in the persistence of unemployment in those countries. This change in the autoregressive coefficients is characterized in our model by different values, for the unemployment persistence before and after the financial crisis.

The mean bias-corrected estimator enables us to test whether the recent financial crisis has produced a significant change in the duration of unemployment in the U.S. and other OECD countries. Based on the bootstrap permutation test, we find some evidence that changes in the persistence of unemployment in these countries are related to the financial crisis in the U.S. economy. The null hypothesis of no change in the persistence of unemployment can be rejected at 10% of probability. Since such test is in its very nature conservative (Efron and Tibshirani, 1993), we conclude that this is an important indication of statistically significant interdependence and collateral effects of the U.S. economy on other OECD countries.

However, our results show heterogeneity in the other OECD countries' response to the external shock represented by the 2007-2008 U.S. financial crisis and offer borderline evidence of an increase in the pattern of unemployment persistence in the other OECD countries which is correlated with such financial crisis. All in all, we believe that this paper, by providing new evidence on the patterns of unemployment persistence in OECD countries, has contributed to the ongoing research on the determinants of such patterns and, therefore, on the policies that can affect them in a welfare-improving way.

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TABLE 3

Searching for structural break in the path of unemployment – Male, 15-24 years old.

	MADF^{GLS}: $H_0 : Q_t^i \approx I(1),\ i = 1,\ldots,25.$				
Countries	N	**Model I** (*t-statistic/break date*)	$\hat{k}$	**Model II** (*t-statistic/break date*)	$\hat{k}$
1. Australia	59	6.93192 / 2008:01[a]	0	6.90902 / 2008:01[a]	0
2. Austria	59	3.98650 / 2005:02	1	4.23378 / 2005:04[c]	1
3. Belgium	59	3.98650 / 2005:02	1	1.81322 / 2004:01	1
4. Canada	59	1.45397 / 2006:02	0	1.67885 / 2009:01	4
5. Czech Republic	59	2.97754 / 2008:01	1	3.00409 / 2007:04	1
6. Denmark	59	1.71052 / 2006:03	1	1.71902 / 2006:04	1
7. Estonia	59	1.48565 / 2006:01	2	1.71086 / 2004:01	2
8. Finland	59	2.46675 / 2007:03	1	2.44624 / 2007:03	1
9. Greece	59	13.15801 / 2008:02[a]	1	12.96061 / 2008:02[a]	1
10. Hungary	59	5.14879 / 2013:01[a]	0	5.11381 / 2012:02[a]	0
11. Ireland	59	3.42036 / 2012:03	2	3.64860 / 2012:01	2
12. Israel	59	3.19529 / 2003:03	2	3.64588 / 2004:01	2
13. Italy	59	16.62656 / 2008:01[a]	0	16.50250 / 2008:01[a]	0
14. Japan	59	2.24311 / 2011:04	1	2.66353 / 2011:04	1
15. South Korea	59	1.09028 / 2013:02	2	1.93443 / 2012:01	2
16. Netherlands	59	2.56271 / 2008:04	0	2.59023 / 2008:02	2
17. New Zealand	59	4.26561 / 2004:01[c]	1	4.19641 / 2005:01[c]	1
18. Norway	59	1.19783 / 2007:04	1	1.54206 / 2003:04	1
19. Poland	59	3.99866 / 2008:04[c]	1	4.12288 / 2008:03[c]	1
20. Portugal	59	4.60661 / 2008:02[b]	1	4.59003 / 2008:02[b]	1
21. Slovak Republic	59	7.30166 / 2007:04[a]	5	7.27610 / 2007:04[a]	0
22. Slovenia	59	3.85016 / 2008:02	2	3.84276 / 2008:02	2
23. Spain	59	12.24767 / 2006:04[a]	4	12.14711 / 2007:01[a]	4
24. United Kingdom	59	4.74697 / 2013:02[b]	0	4.75597 / 2013:02[b]	0
25. United States	59	4.33703 / 2011:04[b]	1	5.30679 / 2009:04[a]	0

Note: Critical values for **MADF^{GLS}**: **Model I** = (1%, 5%, 10%) **-4.87**; **-4.33**; **-3.99** and **Model II** = (1%, 5%, 10%) **-4.86**; **-4.48**; **-4.02**. All critical values were extracted from Perron and Rodríguez (2003, p. 11, Tab. 1b) for N=100. (**a**) Significant at 1% probability; (**b**) significant at 5% probability; (**c**) significant at 10% probability. The number of additional lags (k) was obtained by the sequential *t-test* at 10% of significance, as in Table 2 above.

Source: Authors' calculations from the dataset provided by OECD Statistics.

TABLE 4

Searching for structural break in the path of unemployment – Male, 25-54 years old.

	MADF^{GLS}**:** $H_0 : Q_t^i \approx I(1),\ \ i = 1,...,25.$				
Countries	N	**Model I** (*t-statistic*/*break date*)	$\hat{k}$	**Model II** (*t-statistic*/*break date*)	$\hat{k}$
1. Australia	59	7.70876 / 2007:03[a]	0	7.63891 / 2007:03[a]	0
2. Austria	59	2.06379 / 2012:01	1	2.57484 / 2011:03	1
3. Belgium	59	2.23731 / 2012:01	1	2.35705 / 2011:01	1
4. Canada	60	1.70468 / 2010:03	0	3.26064 / 2009:01	1
5. Czech Republic	59	3.93716 / 2008:01	0	3.93253 / 2007:04	0
6. Denmark	59	2.88385 / 2006:04	0	2.85939 / 2007:01	0
7. Estonia	59	2.87995 / 2006:02	0	2.86291 / 2006:04	0
8. Finland	59	5.32687 / 2007:04[a]	0	5.30487 / 2007:04[a]	0
9. Greece	59	20.73440 / 2008:04[a]	1	20.21569 / 2008:03[a]	1
10. Hungary	59	6.28488 / 2013:01[a]	6	5.60609 / 2013:01[a]	6
11. Ireland	59	3.74790 / 2005:02	1	3.70407 / 2006:02	1
12. Israel	59	2.83504 / 2003:03	1	3.49488 / 2004:01	0
13. Italy	59	18.26625 / 2007:04[a]	0	18.08498 / 2007:04[a]	0
14. Japan	59	2.82871 / 2013:01	0	3.03592 / 2009:02	1
15. South Korea	59	4.04268 / 2003:03[c]	0	4.89534 / 2003:03[a]	1
16. Netherlands	59	6.24444 / 2009:01[a]	1	6.39976 / 2008:03[a]	1
17. New Zealand	59	6.04737 / 2004:04[a]	1	6.00850 / 2006:01[a]	1
18. Norway	59	2.69042 2003:03	1	2.87676 / 2005:01	1
19. Poland	59	3.89831 / 2003:03	2	4.46479 / 2004:02[c]	1
20. Portugal	59	3.65482 / 2008:01	1	3.88144 / 2013:01	0
21. Slovak Republic	59	7.19416 / 2008:03[a]	1	7.16130 / 2007:04[a]	1
22. Slovenia	59	7.29290 / 2008:02[a]	0	7.22615 / 2007:04[a]	0
23. Spain	59	10.52195 / 2006:03[a]	1	10.42061 / 2006:04[a]	1
24. United Kingdom	59	3.52899 / 2004:03	1	3.54116 / 2005:02	1
25. United States	60	5.25157 / 2011:03[a]	1	6.11576 / 2009:01[a]	1

Note: Critical values for **MADF**^{GLS}: **Model I** = (1%, 5%, 10%) **-4.87**; **-4.33**; **-3.99** and **Model II** = (1%, 5%, 10%) **-4.86**; **-4.48**; **-4.02**. All critical values were extracted from Perron and Rodríguez (2003, p. 11, Tab. 1b) for N=100. (**a**) Significant at 1% probability; (**b**) significant at 5% probability; (**c**) significant at 10% probability. The number of additional lags (k) was obtained by the sequential *t-test* at 10% of significance, as in Table 2 above.
Source: Authors' calculations from the dataset provided by OECD Statistics.

TABLE 5

Searching for structural break in the path of unemployment – Female: 15-24 years old.

	MADF$^{\mathbf{GLS}}$**:** $H_0 : Q_t^i \approx I(1),\ i = 1,\ldots,25.$				
Countries	N	**Model I** (*t-statistic*/*break date*)	$\hat{k}$	**Model II** (*t-statistic*/*break date*)	$\hat{k}$
1. Australia	59	6.96369 / 2008:01[a]	0	7.11395 / 2008:01[a]	0
2. Austria	59	2.96130 / 2005:02	0	2.97421 / 2005:04	0
3. Belgium	59	1.26281 / 2004:04	2	1.54774 / 2013:02	2
4. Canada	60	3.39812 / 2007:01	1	3.36043 / 2007:02	1
5. Czech Republic	59	3.19906 / 2008:02	0	3.25223 / 2008:02	0
6. Denmark	59	1.46048 / 2006:04	1	1.78671 / 2012:04	1
7. Estonia	59	1.37804 / 2007:01	4	1.61836 / 2013:01	4
8. Finland	59	2.12120 / 2007:03	2	2.10611 / 2007:03	2
9. Greece	59	10.74996 / 2008:03[a]	0	10.78844 / 2008:02[a]	0
10. Hungary	59	6.74394 / 2013:02[a]	0	6.90984 / 2013:01[a]	1
11. Ireland	59	6.35619 / 2006:03[a]	2	6.29663 / 2007:01[a]	2
12. Israel	59	5.67090 / 2003:03[a]	8	5.69701 / 2003:04[a]	8
13. Italy	59	12.23695 / 2008:01[a]	2	12.12716 / 2007:02[a]	2
14. Japan	59	2.44421 / 2012:01	2	2.80380 / 2010:02	2
15. South Korea	59	1.53104 / 2007:04	0	1.82693 / 2003:03	0
16. Netherlands	59	3.91135 / 2009:01	0	4.16826 / 2008:03[c]	0
17. New Zealand	59	3.14302 / 2004:04	1	3.15381 / 2006:01	1
18. Norway	59	1.05965 / 2008:01	2	1.70110 / 2012:01	1
19. Poland	59	4.55049 / 2009:01[b]	3	4.65906 / 2008:03[b]	0
20. Portugal	59	4.27104 / 2009:02[c]	0	4.24344 / 2008:02[c]	0
21. Slovak Republic	59	5.98762 / 2008:01[a]	0	5.96542 / 2007:02[a]	0
22. Slovenia	59	4.24759 / 2010:01[c]	1	4.31798 / 2008:03[c]	1
23. Spain	59	16.72922 / 2007:01[a]	0	16.61015 / 2007:01[a]	0
24. United Kingdom	59	4.52502 / 2013:02[b]	0	4.33510 / 2013:02[c]	0
25. United States	60	4.14864 / 2013:01[c]	0	4.38740 / 2010:04[c]	0

Note: Critical values for **MADF**$^{\mathbf{GLS}}$: **Model I** = (1%, 5%, 10%) **-4.87**; **-4.33**; **-3.99** and **Model II** = (1%, 5%, 10%) **-4.86**; **-4.48**; **-4.02**. All critical values were extracted from Perron and Rodríguez (2003, p. 11, Tab. 1b) for N=100. (**a**) Significant at 1% probability; (**b**) significant at 5% probability; (**c**) significant at 10% probability. The number of additional lags (*k*) was obtained by the sequential *t-test* at 10% of significance, as in Table 2 above.

Source: Authors' calculations from the dataset provided by OECD Statistics.

TABLE 6

Searching for structural break in the path of unemployment – Results – Female, 25-54 years old.

	MADF^{GLS} : $H_0 : Q_t^i \approx I(1),\ i = 1,...,25.$				
Countries	N	**Model I** (*t-statistic*/*break date*)	$\hat{k}$	**Model II** (*t-statistic*/*break date*)	$\hat{k}$
1. Australia		6.14937 / 2007:04[a]	1	6.09518 / 2007:03[a]	1
2. Austria		3.02842 / 2004:04	1	3.06798 / 2006:03	1
3. Belgium		1.71661 / 2004:03	1	1.96575 / 2004:04	1
4. Canada		1.97362 / 2007:03	0	1.95548 / 2007:03	0
5. Czech Republic		4.09556 / 2008:02[c]	0	4.25695 / 2008:02[c]	0
6. Denmark		5.58042 / 2008:01[a]	4	5.69579 / 2008:01[a]	4
7. Estonia		5.45408 / 2006:01[a]	0	5.39895 / 2006:02[a]	0
8. Finland		5.47826 / 2008:01[a]	1	5.51341 / 2007:04[a]	1
9. Greece		19.74046 / 2009:01[a]	1	19.21147 / 2008:04[a]	1
10. Hungary		8.54099 / 2012:04[a]	0	7.80516 / 2011:02[a]	0
11. Ireland		6.34360 / 2005:03[a]	0	6.20069 / 2006:02[a]	0
12. Israel		3.68587 / 2003:03	1	4.26084 / 2004:02[c]	1
13. Italy		15.85761 / 2007:04[a]	0	15.69683 / 2007:04[a]	0
14. Japan		2.42202 / 2010:04	0	3.05273 / 2009:03	0
15. South Korea		2.73266 / 2013:02	1	2.85079 / 2003:03	0
16. Netherlands		5.03931 / 2011:02[a]	0	4.91496 / 2009:01[a]	0
17. New Zealand		8.04513 / 2006:02[a]	1	7.97601 / 2006:04[a]	1
18. Norway		2.53723 / 2012:03	1	2.52974 2012:02	1
19. Poland		3.59045 / 2009:03	2	3.86205 / 2008:04	1
20. Portugal		3.04150 / 2013:02	0	4.60111 / 2012:04[b]	0
21. Slovak Republic		3.40678 / 2008:04	4	3.68673 / 2008:03	1
22. Slovenia		7.59142 / 2009:01[a]	1	7.93529 / 2008:03[a]	1
23. Spain		15.01973 / 2007:01[a]	1	14.88553 / 2007:01[a]	1
24. United Kingdom		5.32771 / 2004:01[a]	1	5.53172 / 2005:04[a]	1
25. United States		5.31278 / 2012:02[a]	1	5.33547 / 2011:02[a]	1

Note: Critical values for **MADF^{GLS}**: **Model I** = (1%, 5%, 10%) **-4.87**; **-4.33**; **-3.99** and **Model II** = (1%, 5%, 10%) **-4.86**; **-4.48**; **-4.02**. All critical values were extracted from Perron and Rodríguez (2003, p. 11, Tab. 1b) for N=100. (**a**) Significant at 1% probability; (**b**) significant at 5% probability; (**c**) significant at 10% probability. The number of additional lags (k) was obtained by the sequential *t-test* at 10% of significance, as in Table 2 above.
Source: Authors' calculations from the dataset provided by OECD Statistics.

TABLE 8

Bootstrap mean bias-corrected estimates for unemployment in OECD countries – Male: 1 and 2.

Countries (**1**)	$\hat{\alpha}^{(1)}$	HL(1)	$p(1)$	Countries (**2**)	$\hat{\alpha}^{(2)}$	HL	$p(2)$
1. Australia	1.00	∞	1	1. Australia	1.00	∞	2
2. Austria	0.83	3.7	3	2. Finland	1.00	∞	2
3. Greece	0.99	69.0	3	3. Greece	0.97	22.8	4
4. Hungary	1.00	∞	1	4. Hungary	0.91	7.3	4
5. Italy	1.00	∞	1	5. Italy	1.00	∞	1
6. New Zealand	1.00	∞	2	6. South Korea	0.68	1.8	5
7. Poland	1.00	∞	2	7. Netherlands	0.95	13.5	3
8. Portugal	0.96	17.0	4	8. New Zealand	0.98	34.3	4
9. Czech Republic	0.98	34.3	5	9. Slovak Republic	1.00	∞	2
10. Spain	1.00	∞	5	10. Slovenia	1.00	∞	1
11. United Kingdom	0.96	17.0	2	11. Spain	0.99	69.0	2
12. United States	0.93	9.6	4	12. United States	0.98	34.3	2

Note: The maximal order considered was 5 in all models and the BIC criterion was used to select the best lag length. In all cases we specify the model with drift and trend performing 1,000 bootstrap replications. HL is measured in the same time periods.
Source: Authors' calculations from the dataset provided by OCDE Statistics.

TABLE 9

Bootstrap mean bias-corrected estimates for unemployment in OECD countries – Female: 3 and 4.

Countries (1)	$\hat{\alpha}^{(3)}$	HL(1)	p	Countries (2)	$\hat{\alpha}^{(4)}$	HL(2)	p
1.Australia	1.00	∞	1	1. Australia	1.00	∞	3
2. Greece	0.95	13.5	5	2. Czech Republic	0.90	6.6	3
3. Hungary	1.00	∞	1	3. Denmark	0.95	13.5	4
4. Ireland	1.00	∞	3	4. Estonia	0.94	11.2	4
5. Israel	0.83	3.7	2	5. Finland	0.95	13.5	2
6. Poland	0.97	22.8	4	6. Greece	0.99	69.0	3
7. Portugal	1.00	∞	1	7. Hungary	1.00	∞	3
8. Slovak Republic	1.00	∞	1	8. Ireland	1.00	∞	3
9. Slovenia	0.89	5.9	1	9. Israel	0.85	4.3	1
10. Spain	1.00	∞	5	10. Italy	1.00	∞	1
11. United Kingdom	1.00	∞	2	11. Netherlands	0.99	69.0	1
12. United States	1.00	∞	2	12. New Zealand	0.99	69.0	1
----------------------	-----	-----	------	13. Portugal	0.80	3.1	3
----------------------	-----	-----	------	14. Slovenia	1.00	∞	2
----------------------	-----	-----	------	15. Spain	1.00	∞	5
----------------------	-----	-----	-----	16. United Kingdom	0.98	34.3	4
----------------------	------	-----	------	17. United States	0.99	69.0	5

Note: The maximal order considered was 5 in all models and the BIC criterion was used to select the best lag length. In all cases we specify the model with drift and trend performing 1,000 bootstrap replications. HL is measured in the same time periods.
Source: Authors' calculations from the dataset provided by OCDE Statistics.

Appendix

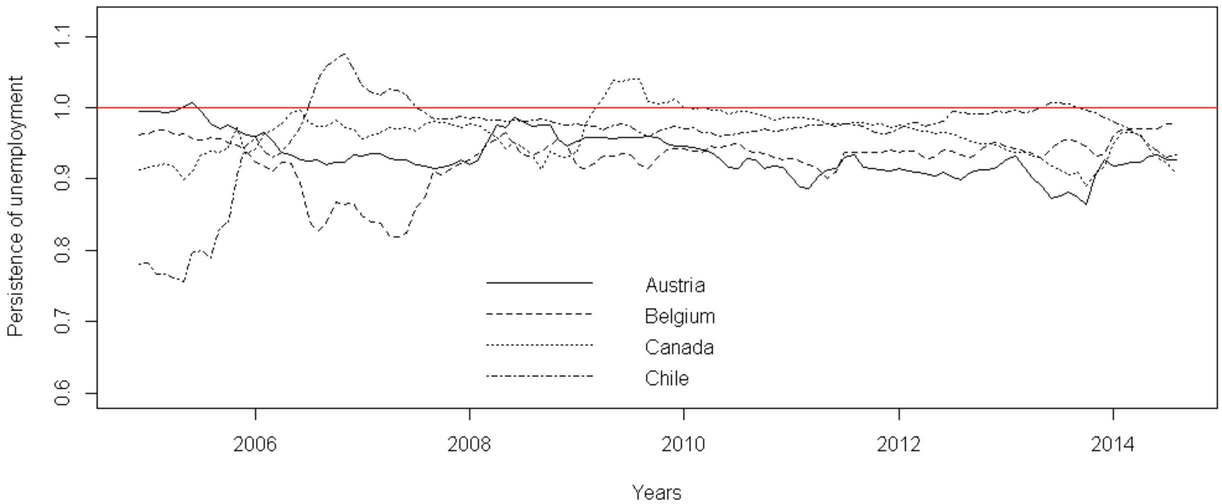


Figure 1a: Path of persistence since 2000, January. Results of moving window for 5 years (*n*=60).

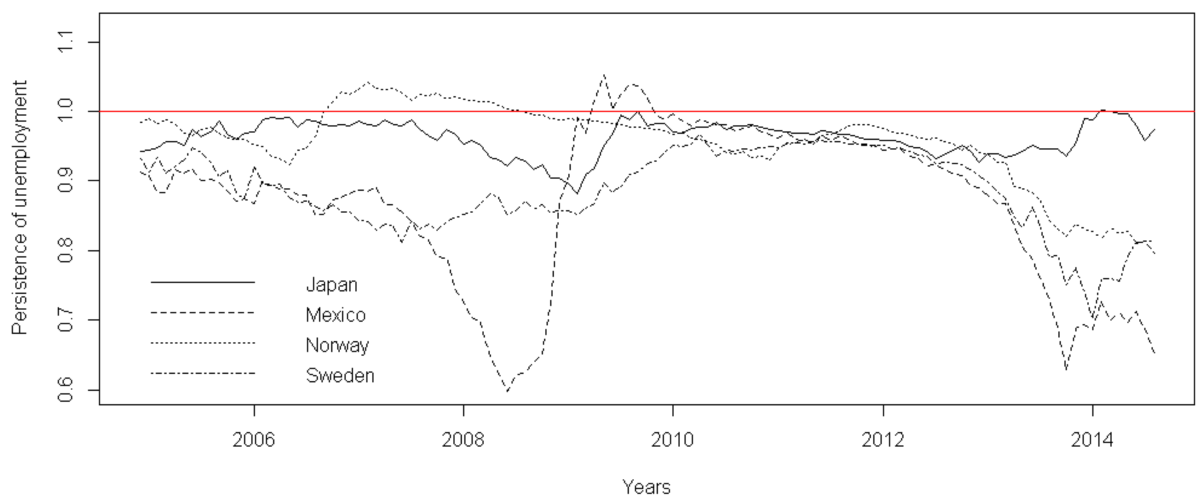


Figure 1b: Path of persistence since 2000, January. Results of moving window for 5 years (*n*=60).

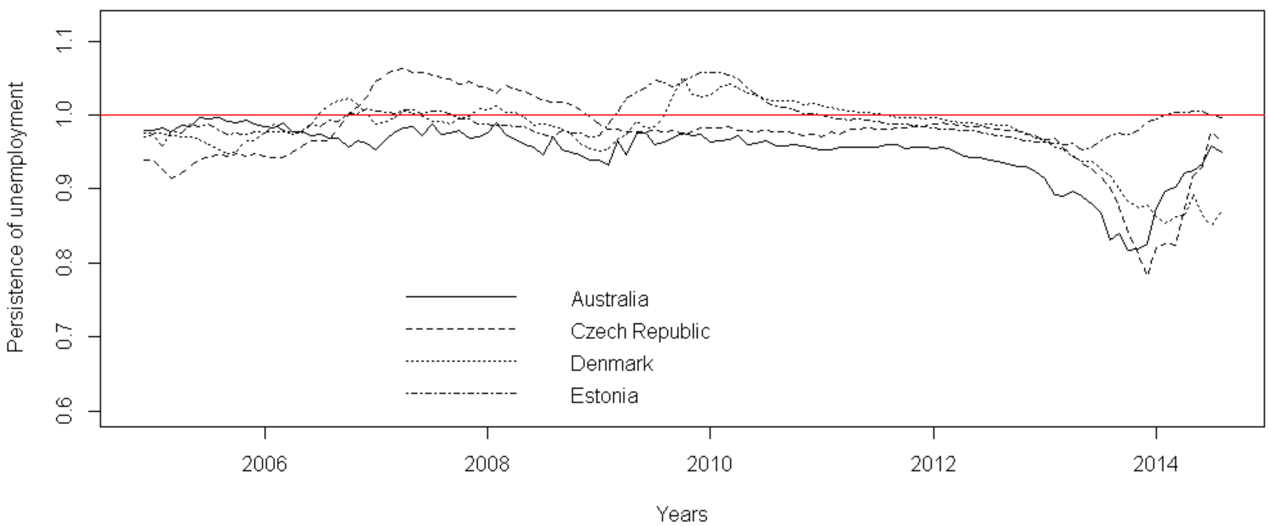


Figure 1c: Path of persistence since 2000, January. Results of moving window for 5 years (*n*=60).

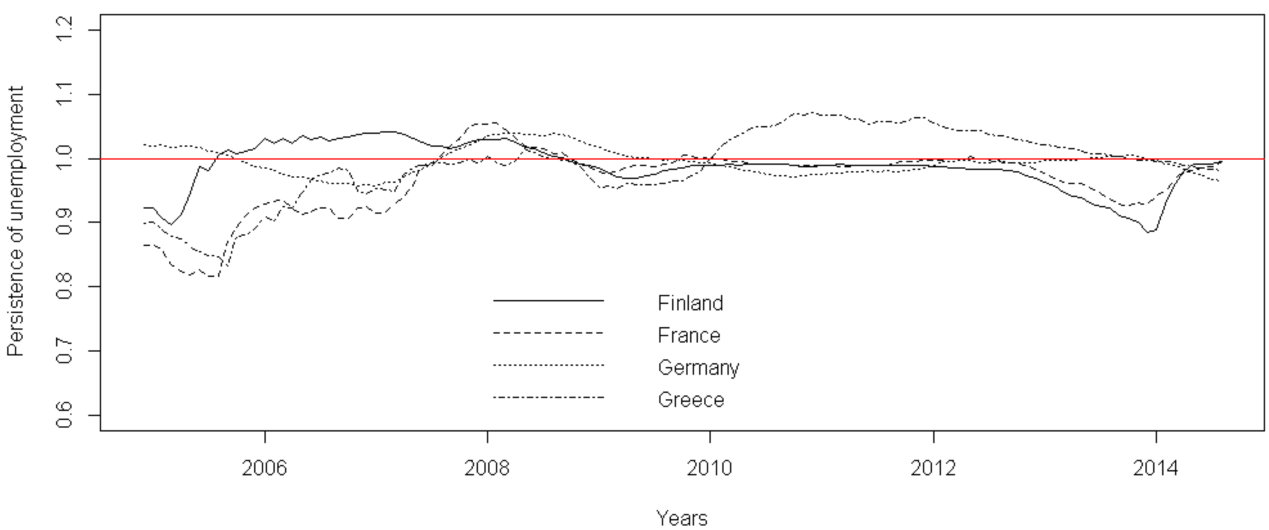


Figure 1d: Path of persistence since 2000, January. Results of moving window for 5 years (*n*=60).

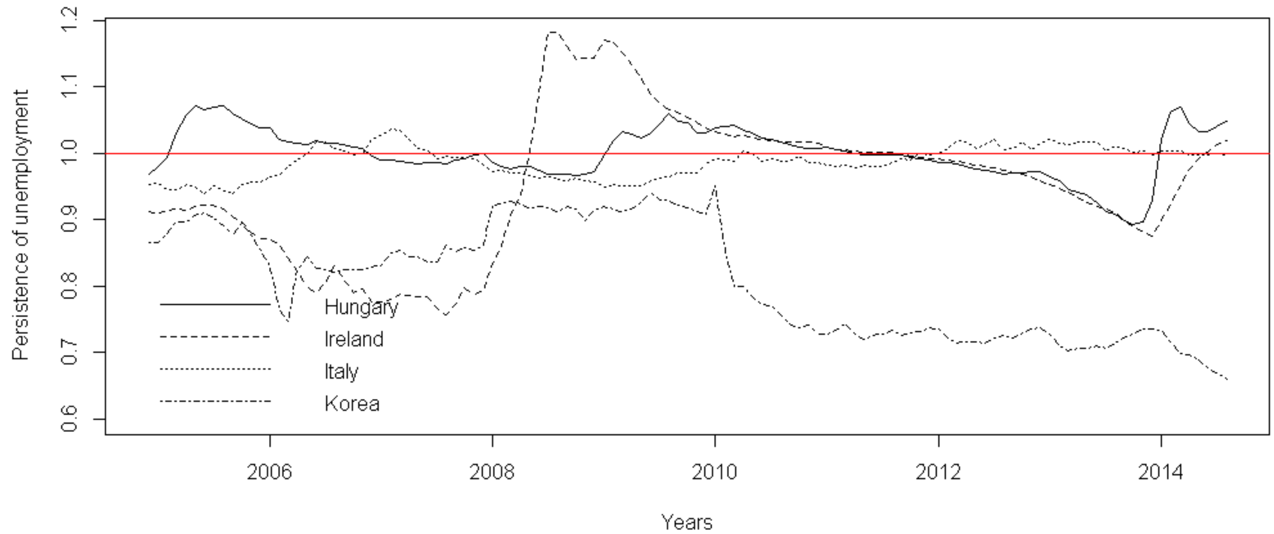


Figure 1e: Path of persistence since 2000, January. Results of moving window for 5 years (*n*=60).

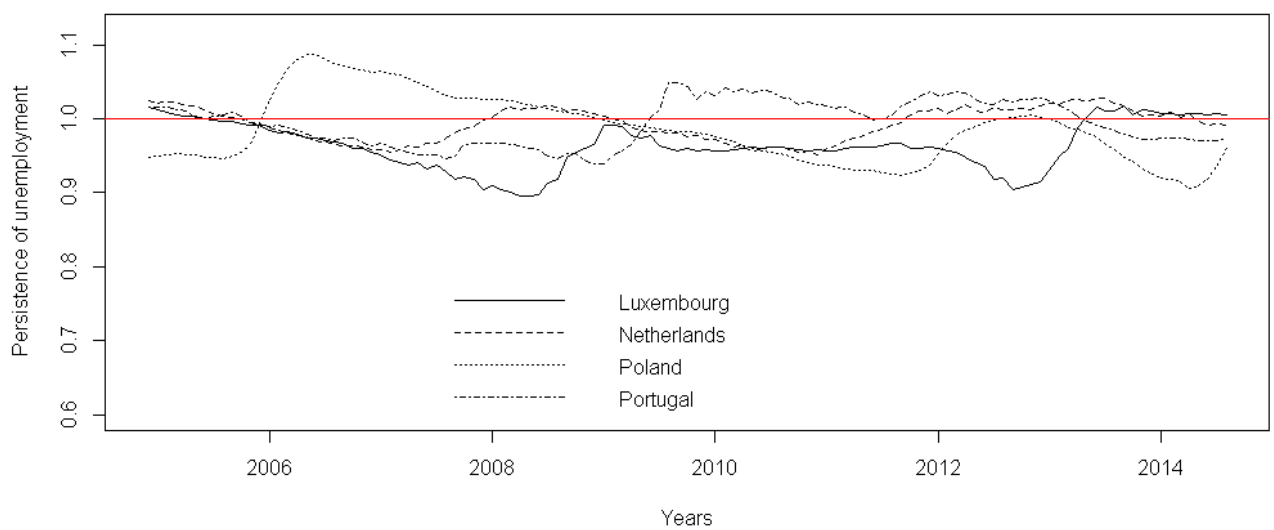


Figure 1f: Path of persistence since 2000, January. Results of moving window for 5 years (*n*=60).

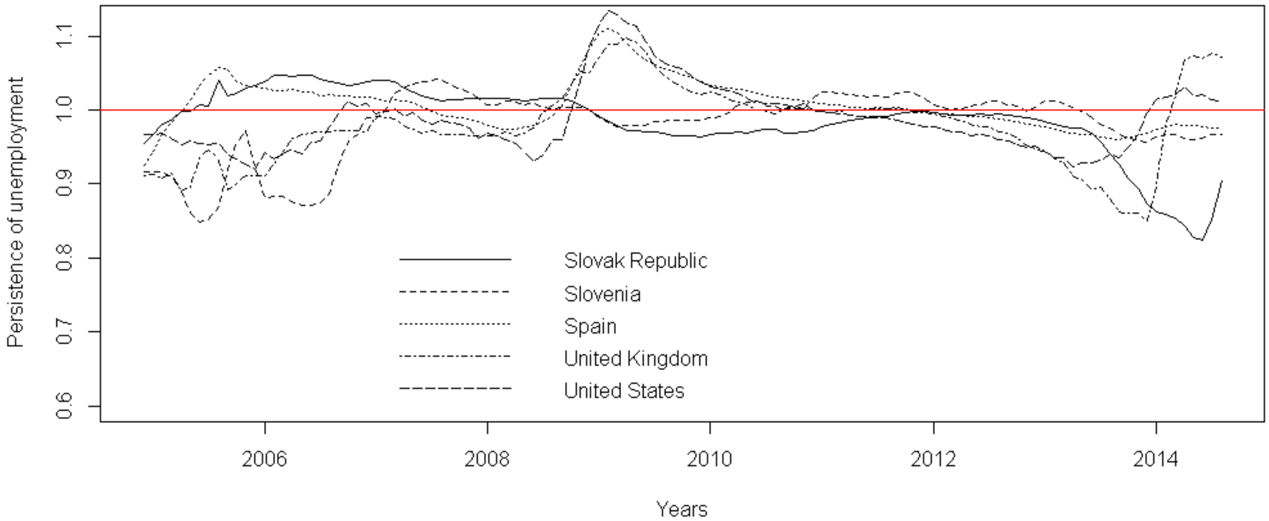
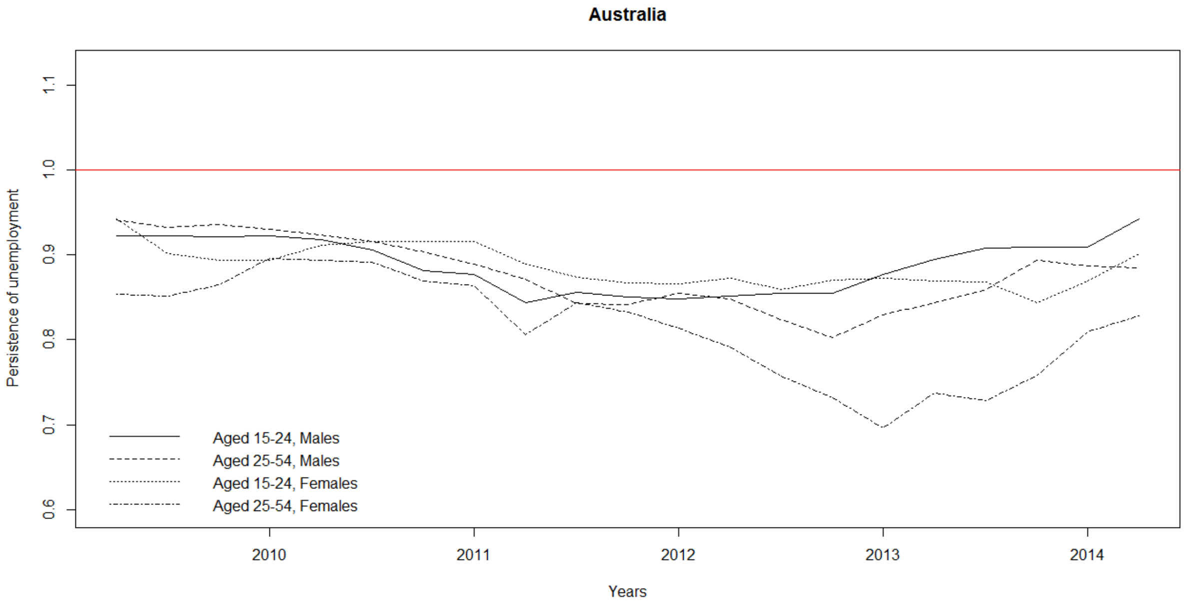


Figure 1g: Path of persistence since 2000, January. Results of moving window for 5 years (*n*=60).



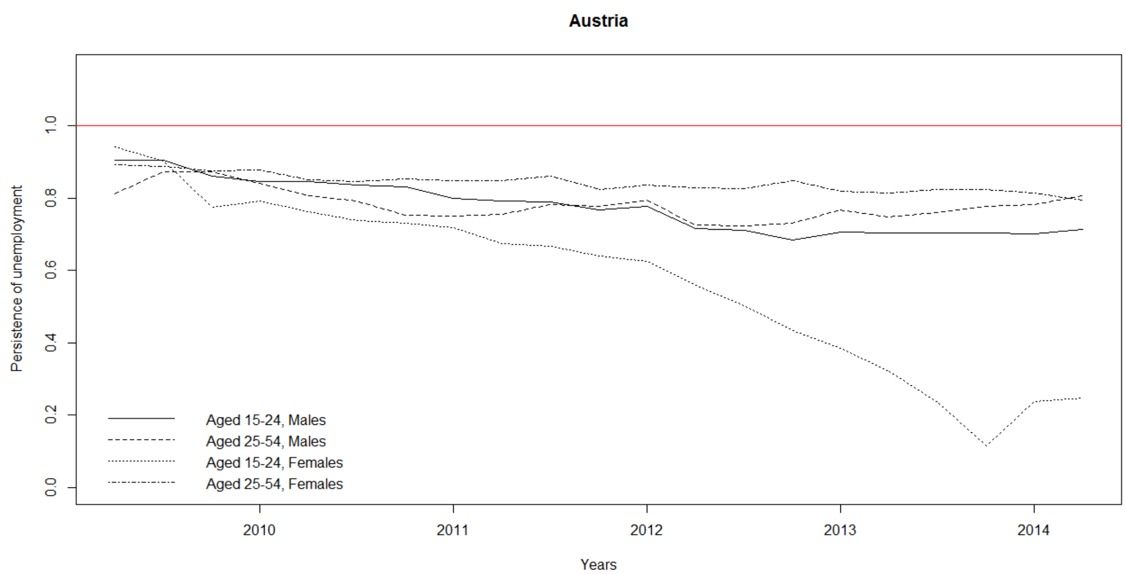
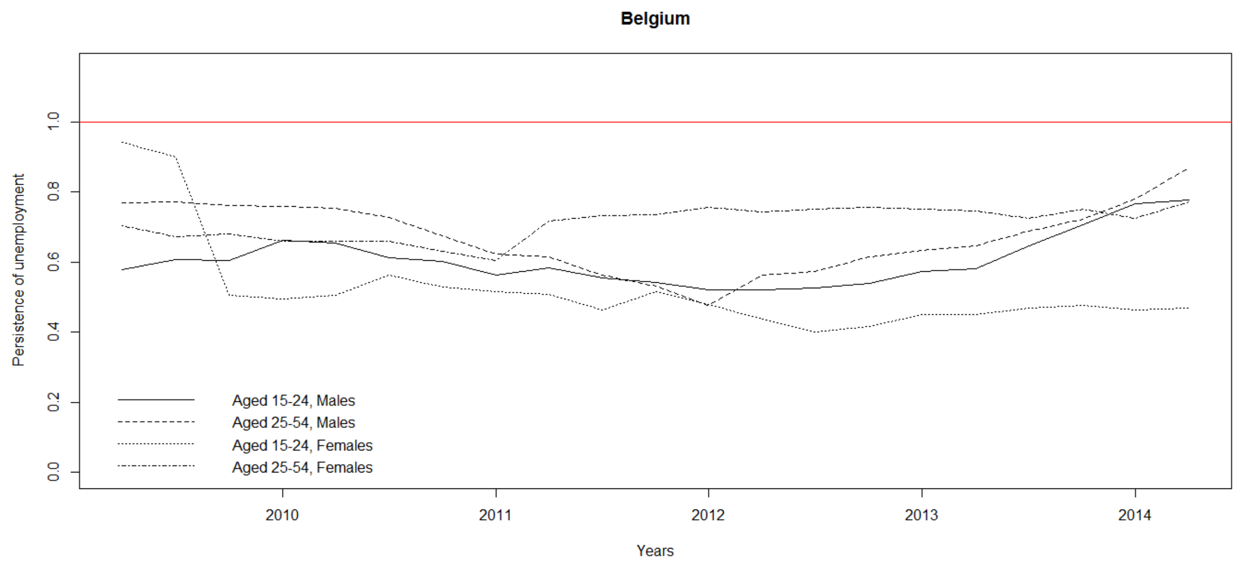
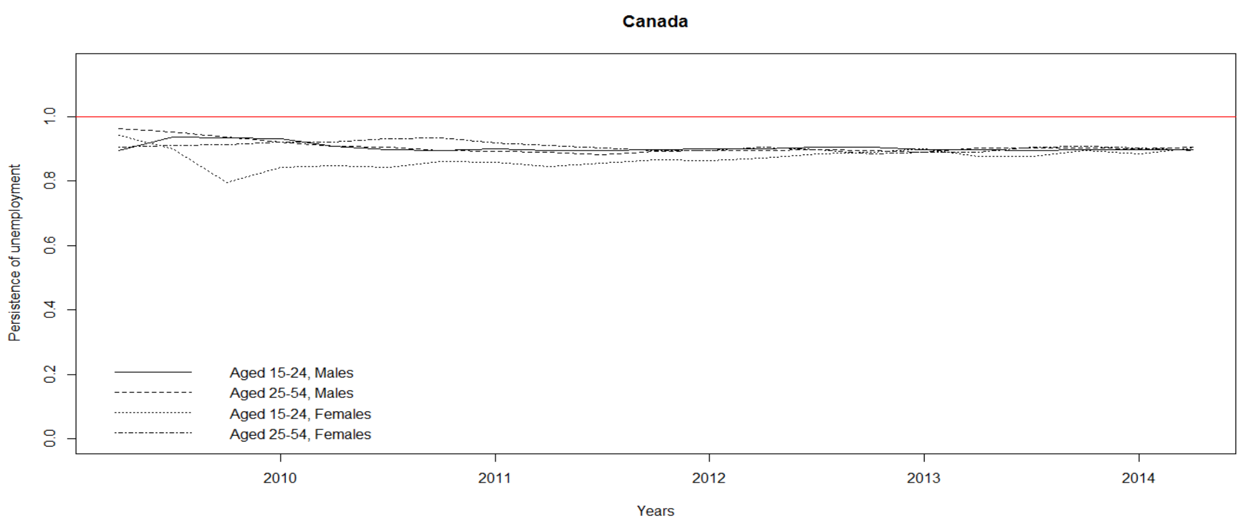


Figure 2a: Path of persistence since 2000QI by gender and age. Results of moving window for 38 periods.


Belgium
Persistence of unemployment
0.0
0.2
0.4
0.6
0.8
1.0
Aged 15-24, Males
Aged 25-54, Males
Aged 15-24, Females
Aged 25-54, Females
2010
2011
2012
2013
2014
Years


Canada
Persistence of unemployment
0.0
0.2
0.4
0.6
0.8
1.0
Aged 15-24, Males
Aged 25-54, Males
Aged 15-24, Females
Aged 25-54, Females
2010
2011
2012
2013
2014
Years

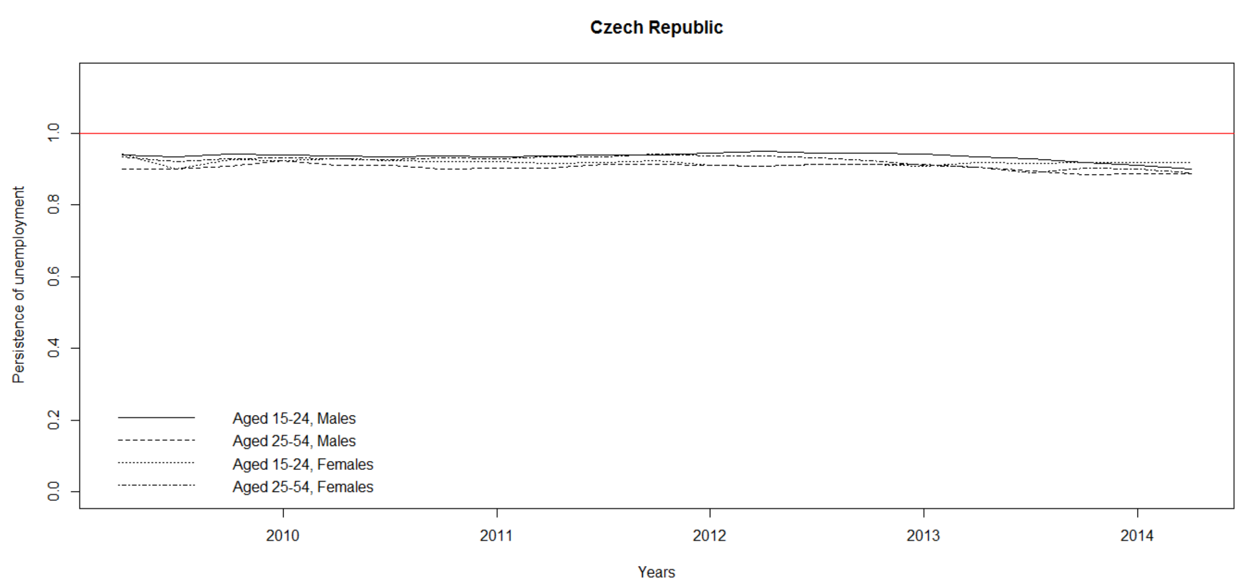
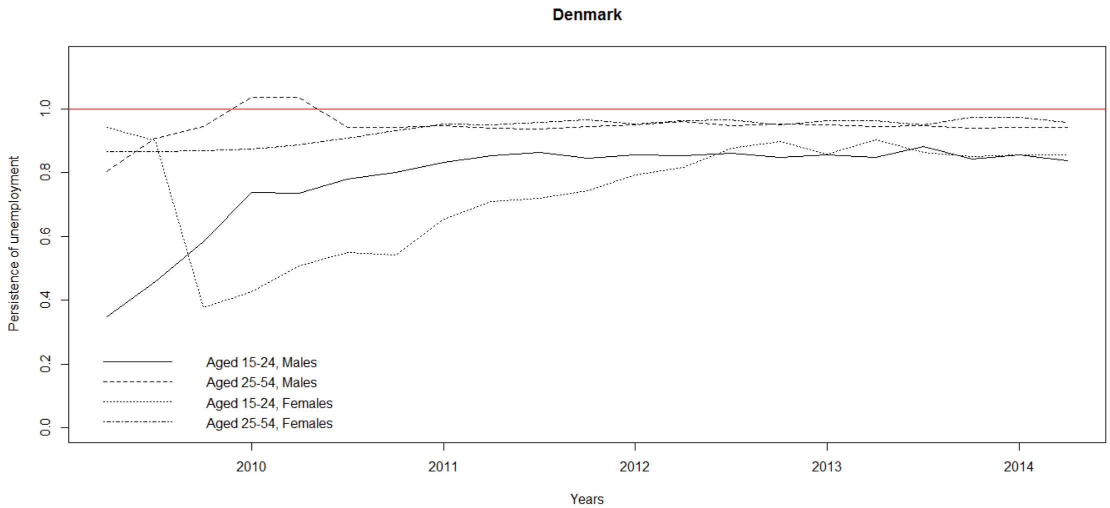
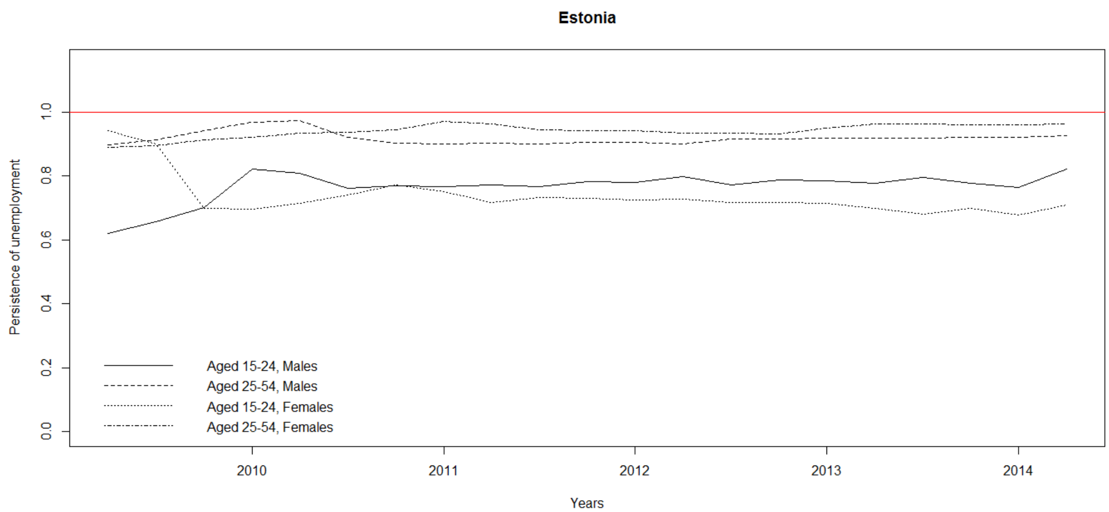

Czech Republic
Persistence of unemployment
0.0
0.2
0.4
0.6
0.8
1.0
Aged 15-24, Males
Aged 25-54, Males
Aged 15-24, Females
Aged 25-54, Females
2010
2011
2012
2013
2014
Years

Figure 2b: Path of persistence since 2000 Q1 by gender and age. Results of moving window for 38 periods.





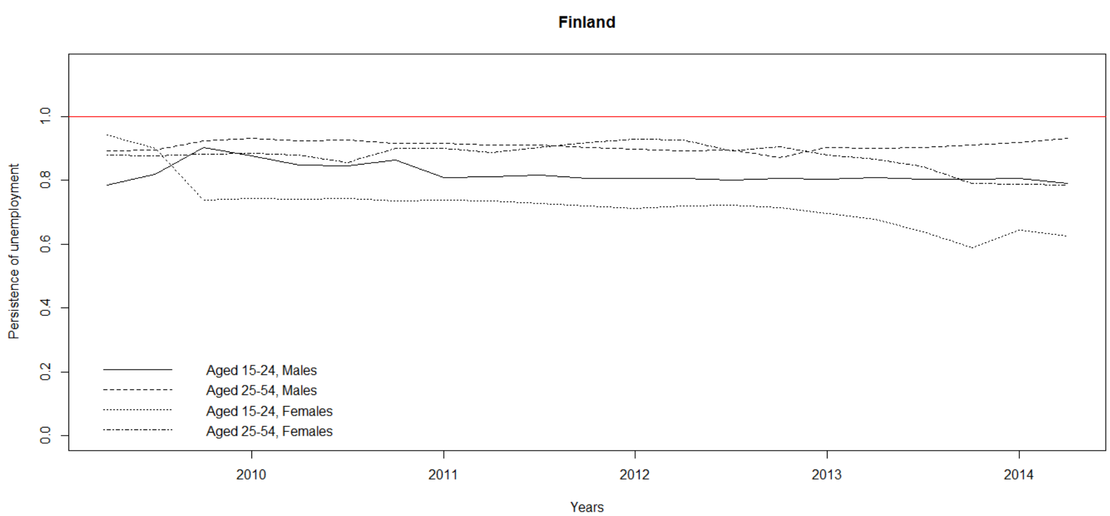


Figure 2c: Path of persistence since 2000QI by gender and age. Results of moving window for 38 periods.

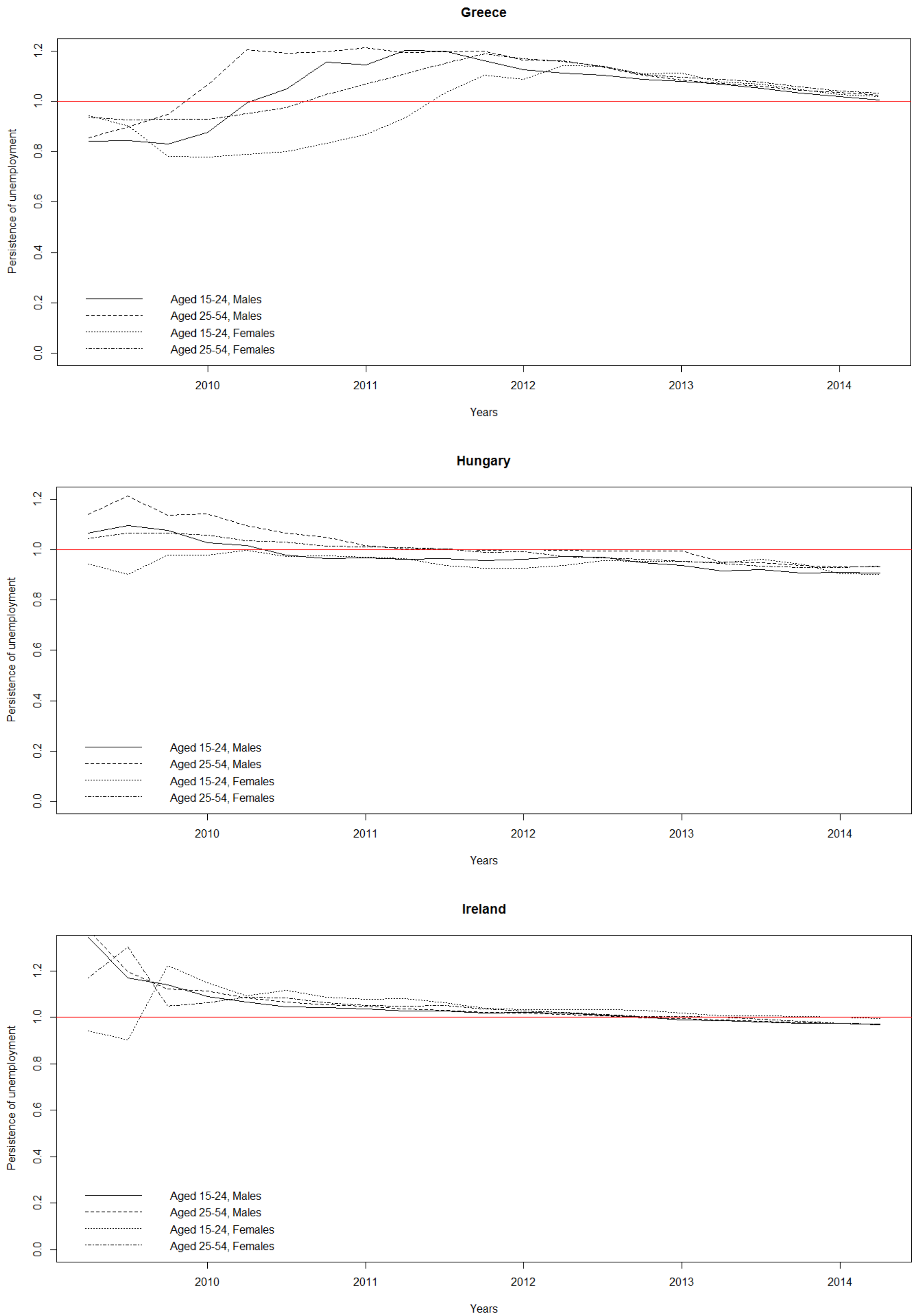


Figure 2d: Path of persistence since 2000 Q1 by gender and age. Results of moving window for 38 periods.

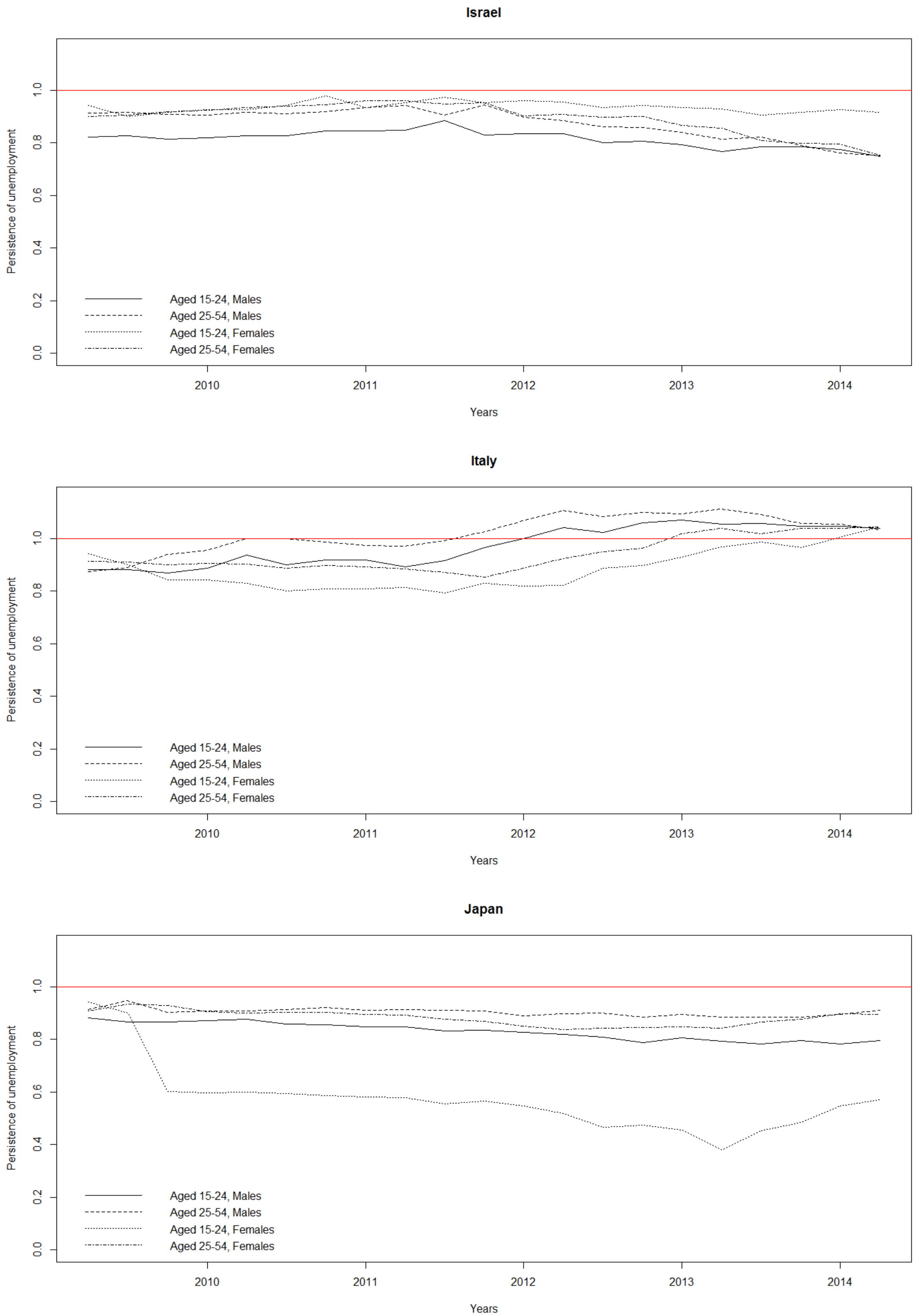


Figure 2e: Path of persistence since 2000 Q1 by gender and age. Results of moving window for 38 periods.

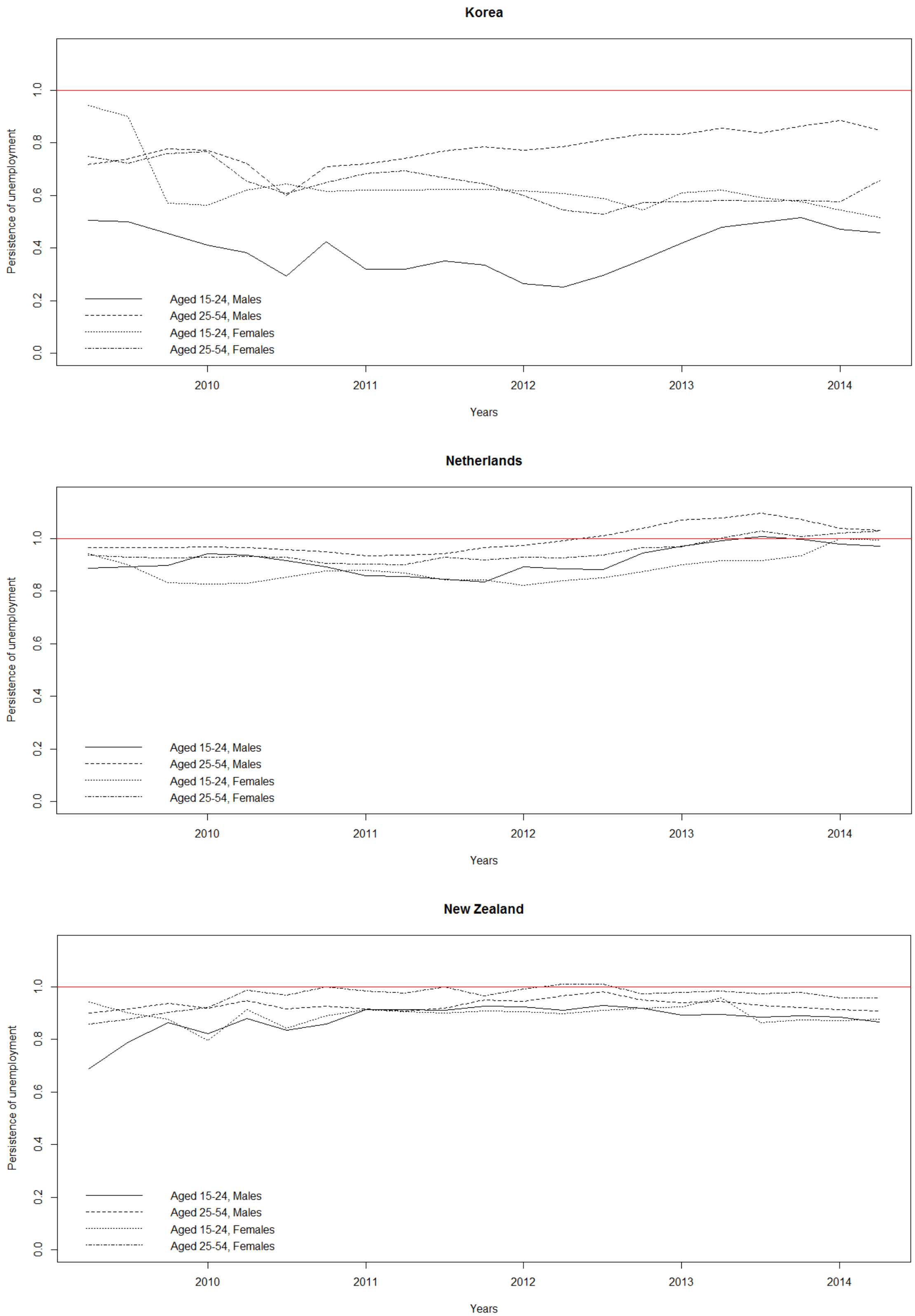


Figure 2f: Path of persistence since 2000 Q1 by gender and age. Results of moving window for 38 periods.

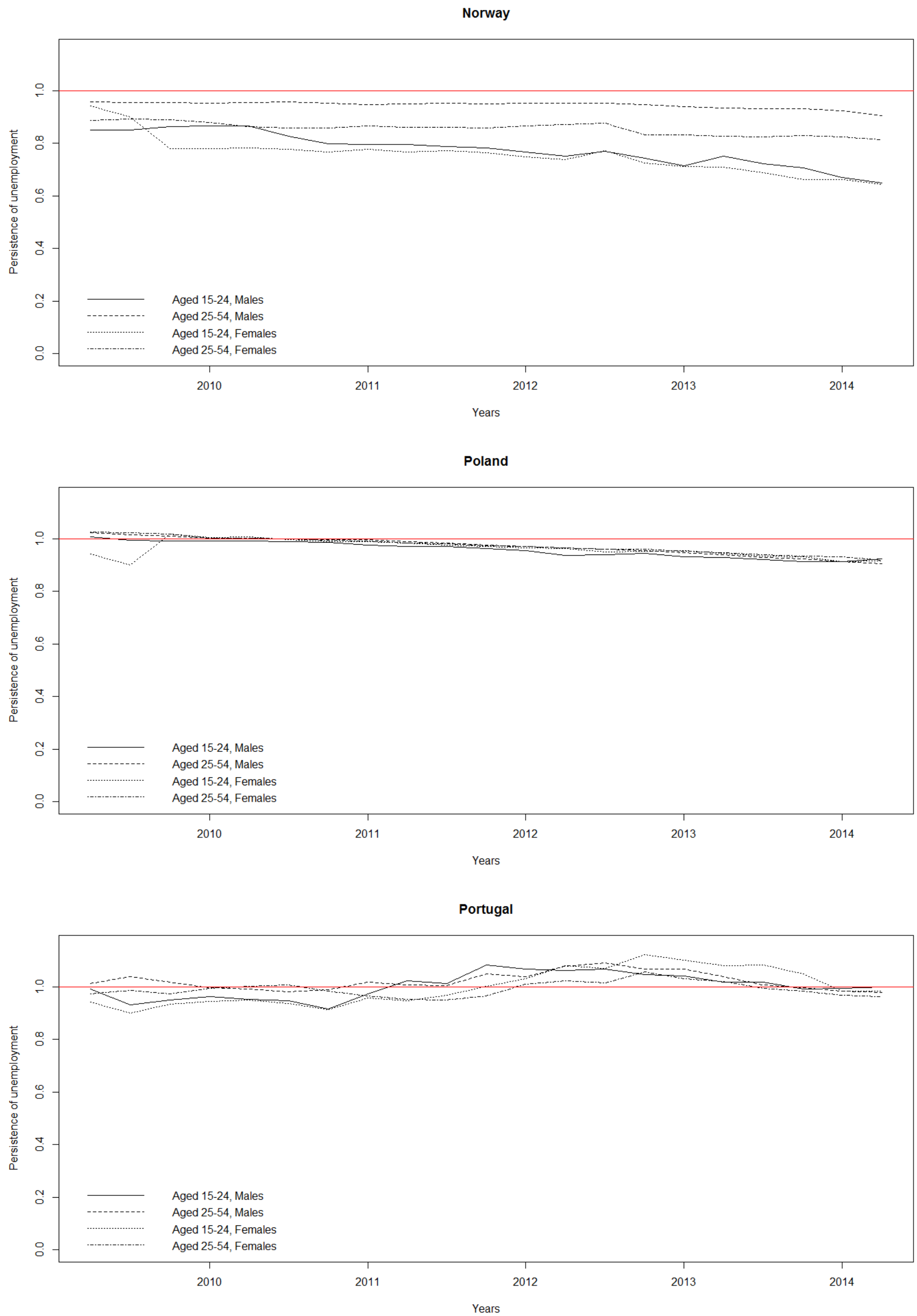


Figure 2g: Path of persistence since 2000 Q1 by gender and age. Results of moving window for 38 periods.

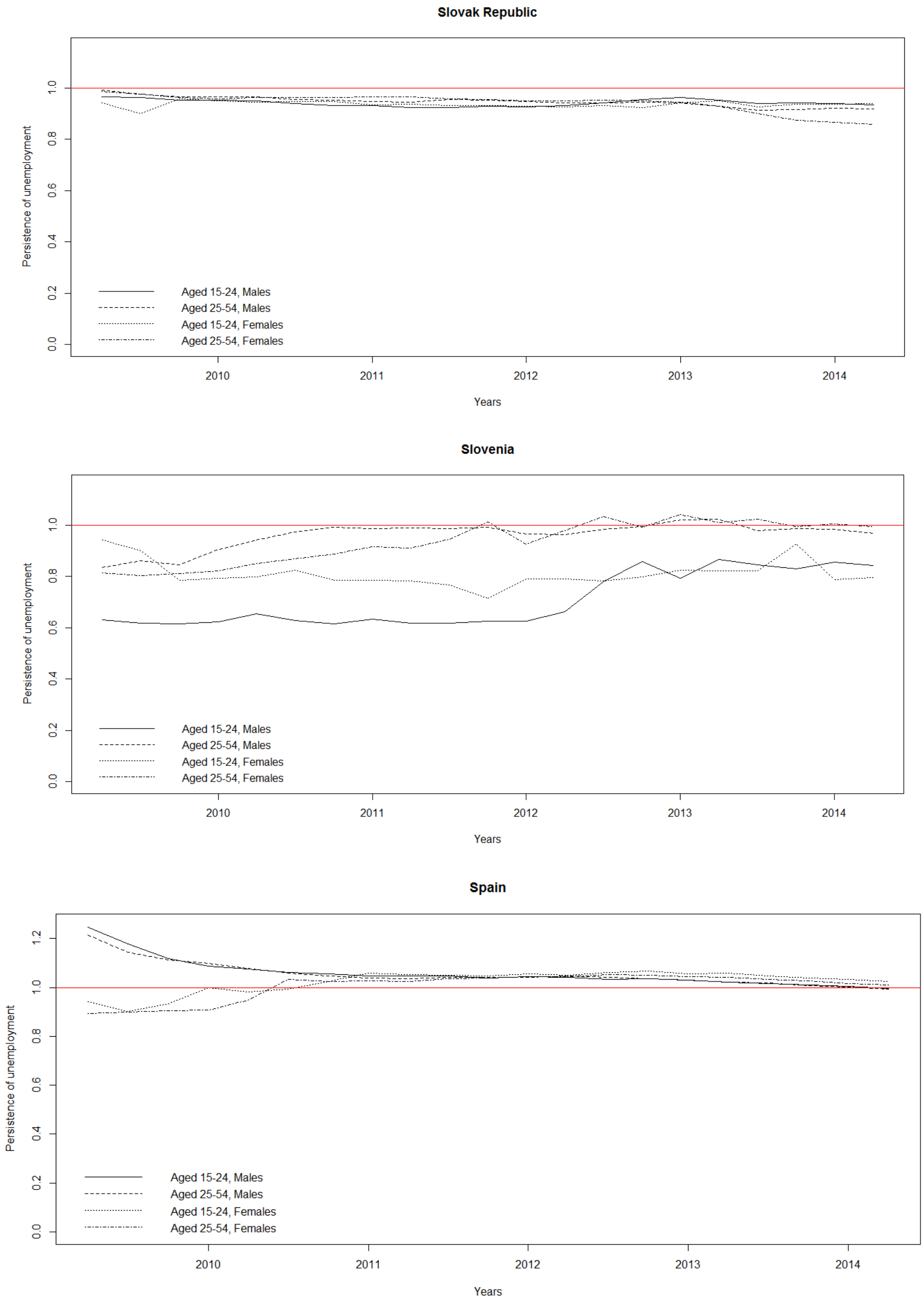
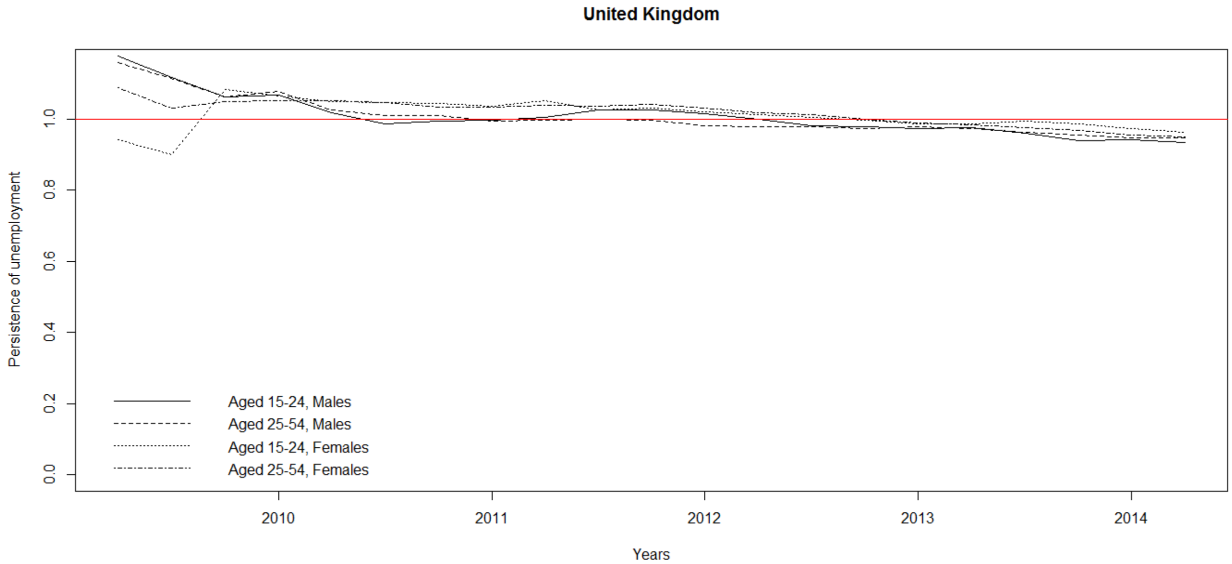


Figure 2h: Path of persistence since 2000 Q1 by gender and age. Results of moving window for 38 periods.



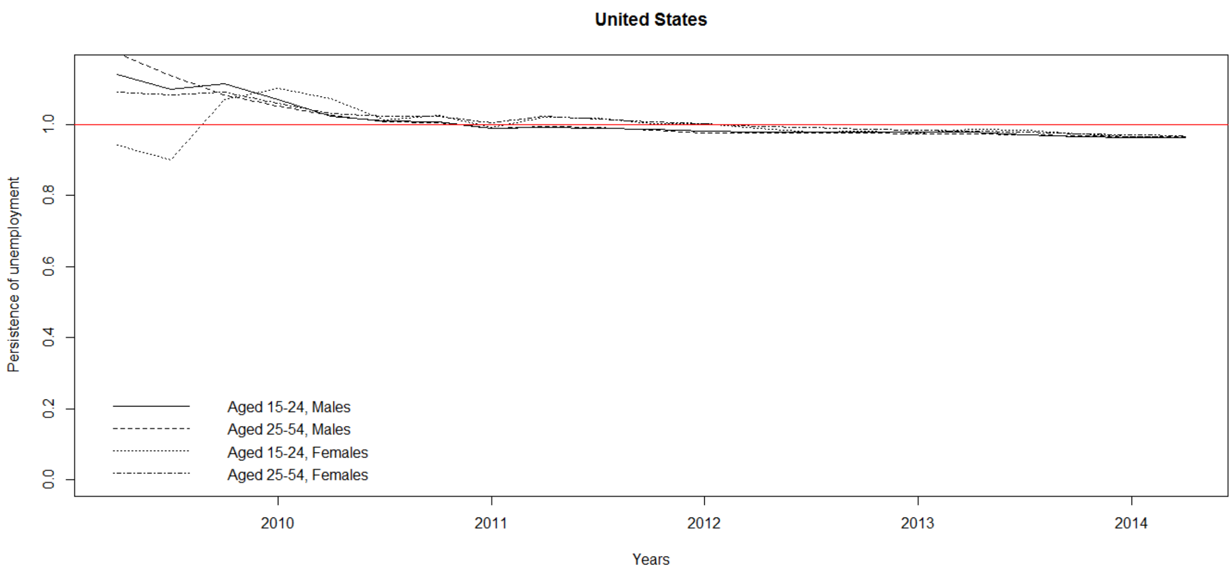


Figure 2i: Path of persistence since 2000 Q1 by gender and age. Results of moving window for 38 periods.