The unemployment structure of the US States.

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Abstract

This paper analyses the time series properties of the unemployment rates of the 50 US States, as well as the global rate of the USA. Our results, based on the use of ADF-type tests, show that the inclusion of some breaks is vital in order to reduce the persistence on these rates. Thus, we can reject the unit root null hypothesis versus a double mean-shifted stationary alternative for 46 States and for the US total rate. We also …nd that the behavior of this latter rate is not congruent with that of the States, implying the presence of some aggregation problems which have not been commonly accounted in the literature.

Keywords: Unemployment rate; Unit Roots; Structural Breaks; Aggregation; NAIRU

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1 Introduction

The analysis of the unemployment rates of the OECD countries has recently attracted a great deal of interest on the part of both theoretical and applied researchers. This interest is easily understood if we bear in mind that unemployment is one of the most important economic problems for these countries. The majority of the papers dedicated to this theme have focused on the study of the time properties of the unemployment rate, given the economic implications involved by the acceptance or rejection of the unit root null hypothesis for this rate. For example, if we accept this hypothesis, we can then interpret this result as evidence in favor of the presence of hysteresis in the evolution of the unemployment rate of a given country. By contrast, its rejection would provide evidence in favor of the so-called structuralist models. The di¤erences between these two lines of thinking are very important. Hysteresis models consider that current unemployment depends on its past values, which implies that all shocks have a permanent in‡uence on the natural rate. Consequently, this rate is changing over time, with its value depending on the size of the shocks. Thus, we cannot use it in order to compare the evolution of the labor markets of a variety of countries, regions or states. By contrast, structuralist models endogeneize the natural unemployment rate, allowing for the presence of some infrequent permanent shocks, which change the value of the equilibrium unemployment rate. Therefore, it is obvious that this concept is meaningless from an hysteresis perspective, but provides extremely useful information under the structuralist approach, in that it admits the existence of an equilibrium value, although this can occasionally change. Thus, provided that we can o¤er evidence in favor of the structuralist hypothesis, then this equilibrium unemployment rate, also called NAIRU, can be considered as an useful tool in order to compare labor markets.

Against this background, it is perfectly understandable that a number of papers have focused their attention on a careful determination of the time properties of the unemployment rates of the OECD countries. In this regard, some works initially provided clear evidence of the existence of a signi…cant amount of persistence in the unemployment rates and, consequently, o¤ered support for the idea that those models based on hysteresis could more appropriately represent the behavior of unemployment rates. Here, we can cite the papers of Blanchard and Summers (1986), Jaeger and Parkinson (1994) and Roed (1996). However, some more recent papers have shown that most of the persistence exhibited by these rates was spuriously created by neglecting the presence of some possible structural breaks. Once this e¤ect is detected and discounted, the persistence is substantially reduced and, consequently, the evidence in favor of the hysteresis models is much less, if it exists at all. In this regard, attention should be drawn to Bianchi and Zoega (1998), Arestis and Mariscal (1999), Phelps and Zoega (1998), Nickell (1998) or Papell et al. (2000).

We should nevertheless recognize that a conclusion in favor or against these two models is not always easy to draw and, consequently, the discussion is far from closed. For example, in our view there are two additional factors that, to the best of our knowledge, have not been jointly considered in the literature and that might change the direction of the empirical evidence in favor or against the above-mentioned theoretical models. First, a possible extension of the above mentioned lines of research is the inclusion of multiple structural changes in the evolution of the unemployment rates, as is done in Arestis and Mariscal (1999). The second issue that we should consider is the possible e¤ect of aggregation on the analysis of the unemployment rates, a question which has been partially raised in Song and Wu (1997) and Payne et al. (1999). Whilst we consider that this should not a¤ect the European case to any great extent, in that its regions have a reduced dimension and a more homogeneous pattern of behavior, it may clearly have a greater incidence in the case of the USA. At this point, we should recall that most of the analyses for the USA case are based on the study of the total US unemployment rate. However, this rate is a weighted average of the rates of the di¤erent States, and thus it is quite reasonable to think that whilst the evolution of the unemployment rates of these States should be somehow related, they are far from being identical. Therefore, the analysis based on the total US unemployment rate may lead to distorted conclusions. Reasons such as di¤erences in population or in the economic structure leads us to assume that the economic shocks should not a¤ect all the States is an equivalent manner and, therefore, the evolution of regional unemployment rates could be quite di¤erent. In this regard, we should cite the papers of Partridge and Rickman (1995, 1997), where the US States unemployment rate dramatically depends on factors such as the industrial composition, education or amenities, amongst others. The e¤ect of this heterogeneous behavior can clearly alter the time properties of the total US unemployment rate and, therefore, it seems to sensible to carry out both the analysis of the unemployment rate for the whole US economy as well as for each of the States which compose it.

In the light of the above, the aim of the paper is to show how a study of unemployment at a regional level can be extremely useful when seeking to characterize the unemployment rate structure of a given country. As a case study, we consider that of the US States to be particularly appropriate, given its earlier-mentioned heterogeneity. The rest of the paper is organized as follows. Section 2 is devoted to specifying the econometric models and unit root test. In Section 3 we analyze the time properties of the unemployment rates of the different US States and, for the purpose of comparison, of the total US unemployment rate. This analysis is carried out using the Dickey-Fuller family of tests, which includes those statistics that allow for the presence of some structural breaks in the model specification. We find that the inclusion of these breaks help us to reduce the degree of persistence and, consequently, the evidence in favor of the unit root null hypothesis. In Section 4 we discuss the economic implications of the results from the point of view of the debate between structuralist and hysteresis schools. We also offer an approximation to the estimation of the natural rate of unemployment for all the US States. Section 5 closes the paper with a review of the most important conclusions.

2 Unit root tests and structural breaks

As we have already stated, the goal of this paper is to analyze the time properties of the unemployment rates for the 50 US States. To that end, we have chosen to test for the unit root null hypothesis by way of ADF type tests. The rejection of the unit root null hypothesis leads us to conclude that all the shocks influencing the behavior of the unemployment rate have a transitory nature, and thus these rates are fluctuating around a mean value, with this being interpreted as evidence in favor of the equilibrium or structuralist models. By contrast, if we accept the unit root null hypothesis, this would imply that these shocks are permanent and, therefore, that the structuralist models should be questioned.

Several statistics can be used in order to analyze the integration order of a variable. The most popular method is to study whether the autoregressive parameter is equal to 1 by way of the pseudo t-ratio proposed in Dickey and Fuller (1979). Under the presence of an autocorrelation pattern in the residuals, we can use the modifications developed in Said and Dickey (1984) or carry out the non-parametric corrections proposed in Phillips and Perron (1988). None of these methods offer good properties when the variable being studied exhibits some structural breaks. Perron (1989, 1990) and, more recently, Montañés and Reyes (1998a) have shown that, under the omission of such breaks, the estimator of the autoregressive parameter goes towards 1. Therefore, these statistics can be biased towards the acceptance of the unit root null hypothesis and, if our intuition is that the variable being analyzed can be influenced by the presence of some structural breaks, alternative methods should be considered.

To that end, let us assume that the variable being studied does not exhibit a trend behavior, but can exhibit some changes in the mean across the sample.

If we extend the seminal work of Perron (1990) to the case of multiple breaks, and exclusively considering that case in which the breaks gradually affect the behavior of the variable, we can test for the integration order of y_t as follows. First, we should estimate this model:

$$y_t = \mu + \sum_{i=1}^{n} \delta_i DTB_{it} + \sum_{i=1}^{n} d_i DU_{it} + \rho y_{t-1} + \sum_{i=1}^{k} \phi_i \Delta y_{t-i} + \varepsilon_t \qquad (1)$$

and, subsequently, obtain the pseudo t-ratio for testing whether the autoregressive parameter is equal to 1. In the previous equation, DTB_{it} is a pulse variable that takes the value 1 if $t = TB_i + 1$ and 0 otherwise and DU_{it} takes the value 1 if $t > TB_i$ and 0 otherwise, where TB_i $(i = 1, \ldots, n)$ are the different periods where the mean of the variable changes. For the sake of simplicity, we will assume throughout this paper that $TB_i = \lambda_i T$ and, further, $\lambda_i > \lambda_{i+1}$. The lags of Δy_{t-i} are included in (1) in order to remove the possible presence of autocorrelation in the residuals, with k being commonly referred to as the lag truncation parameter. Under the null hypothesis, the value of the autoregressive parameter should be equal to 1 and, similarly, $d_i = 0$, $\delta_i \neq 0$ $\forall i = 1, 2, \ldots, n$. By contrast, under the stationary alternative hypothesis, it is to be expected that $|\rho| < 1$ and that $d_i \neq 0;\ \forall i = 1, 2, \ldots, n$.

The distribution of the pseudo t-ratio depends on some nuisance parameters, even when the innovations are not autocorrelated. These parameters are the number of breaks (n) and the break fraction parameters (λ_i). Perron (1990) and Clemente et al. (1998) derive the asymptotic distributions of this statistic under the assumption that these parameters are a priori known. The former considers the case $n = 1$, with the latter considering the case $n = 2$.

We can also follow a much more general approach by considering the time of the breaks as being endogenous variables. Several methods have been developed in this regard. For example, it is possible to determine the time of the breaks by minimizing the value of the pseudo t-ratio for all the range of possible values of the parameter λ. Another possibility is to select the time of the break by optimizing the value of the statistic that measures the significance of the magnitudes of the breaks under the alternative hypothesis, and then testing for the unit root null hypothesis by way of its corresponding pseudo t-ratio. This second strategy offers some power improvement, according to the results of various Monte Carlo studies carried out in Perron and Vogelsang (1992).

All these tests are based on the study of the pseudo t-ratio. A different testing procedure is to use a statistic which takes into account not only the value of the estimation of the autoregressive parameter, but also the estimations of the parameters which measure the magnitude of the breaks under the alternative hypothesis. This involves the study of the joint null hypothesis H_o: $\rho = 1$, $d_i = 0 \ \forall i = 1, 2, \ldots, n$ by way of a pseudo F-ratio. If this statistic takes a value close to 0, the variable is better characterized as being integrated; by contrast, if it takes a value far from 0, then the variable is considered as being stationary around some changes in the mean.

Montañés and Reyes (1998b) and Montañés (1999) derive the asymptotic distributions for the cases $n = 1$ and $n = 2$, respectively, and we refer to these statistics as Φ_1 and Φ_2. Once again, these distributions depend on the above-mentioned nuisance parameter. Thus, it is necessary to estimate them or, at least, to assume that they are a priori known in order to be able to apply them. Now, the time of the breaks is determined by maximizing the value of the pseudo F-ratios, obtaining the max Φ_1 and max Φ_2 statistics. The results of some Monte Carlo experiments carried out in these two papers reflect some improvements in their power when compared to the pseudo t-ratios. Furthermore, these statistics have the additional advantage that they explicitly take into account the magnitudes of the breaks under the alternative hypothesis, an aspect which is not considered by the pseudo t-ratios. Given these advantages, we will use them for the purpose of our empirical analysis, considered in the following Section where we present the results of the application of these pseudo F-ratios to the study of the time properties of the unemployment rates in USA.

3 Empirical application: Analysis of the US unemployment structure

In order to analyze the time properties of the unemployment rates in the USA, we have chosen to apply the statistics discussed in the previous Section. We could have followed alternative routes to that end. For example, we could have chosen to determine the possible presence of mean shifts by way of the procedure proposed in Bai and Perron (1998) and, once these breaks have been discounted, then to test for the unit root null hypothesis. Indeed, this is the method used in Papell et al. (2000) for proving that the unemployment rates of the OCDE countries can be characterized as being mean-shifted variables. Another interesting alternative would have been to apply the Markov switching regression models (see Hamilton, 1989) to determine the existence of different regimes in these variables. An example of the application of this procedure to the analysis of unemployment rates can be found in Bianchi and Zoega (1998). However, whilst they are all of interest, we do not feel that any of these alternatives provide an explicit way for testing the unit root null hypothesis under the presence of changes in the mean of the variable being studying, and thus our preference for the ADF-type statistics, appropriately extended in order to allow for the presence of some changes in the mean. In what follows, we will discuss the results obtained from the applications of these statistics.

Our database contains information for the 49 continental US States, plus the District of Columbia. We have also included the data for the total US unemployment rate so as to obtain a global view of the US unemployment structure. The length of the data base has been determined by the availability of data for the US States. Our sample size contains seasonally adjusted quarterly data and it covers the period from 1978:1 to 1999:4. These data were obtained from the US Bureau of Labor Statistics.

Before analyzing the results that we have obtained, we should recall that a similar study has already been carried out in Song and Wu (1997), where these authors do not …nd a signi…cant body of evidence against the presence of a unit root in the unemployment rate of 48 US states. However, there are two main di¤erences between their paper and ours. First, the results presented by these authors are based on the use of annual data for the sample 1962-1993, and thus our results are not directly comparable. Secondly, they include a deterministic trend in the model speci…cation which allows them to obtain the unit root tests . Additionally, they limit the number of breaks to 1. We consider that the inclusion of this deterministic trend is not justi…ed either from a theoretical or, more importantly, for an empirical point of view. Thus, we consider that it is more sensible to test for the unit root null hypothesis without the presence of the deterministic trend. Furthermore, it is possible that the inclusion of a single break may not be enough to capture the evolution of the di¤erent unemployment rates.

Against this background, the most interesting results are presented in Tables 1, 2, and 3, with these Tables reporting the unit root tests for the case of 0, 1 and 2 breaks, respectively. In all these cases, the selection of the lag truncation parameter was carried out by using the so-called k(t) method, recommended in Ng and Perron (1995). This involves a general-to-speci…c strategy based on the analysis of the single signi…cance of the k ¡ th lag of ¢y_t. We assume that kmax = 5. We have focused our analysis on the case where the breaks are supposed to have a gradual e¤ect on the variable being studied, with this being commonly known as the innovational outlier case. We have also obtained the results for the so called additive outlier case, which implies that the break suddenly modi…es the behavior of the variable. However, we have chosen to not report these results, given that we consider the former to be more appropriate in order to capture the evolution of the unemployment rates in the US States. Additionally, they do not substantially modify the results presented in Tables 2-4. When analyzing these results, we will consider that there exists strong evidence against the unit root null hypothesis whenever the di¤erent statistics employed lead us to reject it for a signi…cance level lower or equal to 5%. By contrast, if we can reject this hypothesis for a 10% signi…cance level, then we will consider that there is weak evidence against it.

Let us begin the analysis by considering the results for the case where no breaks are included in the model speci…cation. We can observe from Table 1 that the ADF test provides strong evidence against the presence of a unit root for the unemployment rates of 27 US States, whilst for the rest this evidence is weak, or we simply cannot reject the unit root null hypothesis. We can also accept the presence of a unit root for the total US unemployment rate. In this latter case, it is of interest to note that the value of the half-life measures is 16:98, much higher than the other measures, which never exceed 2:0. Furthermore, the value of the half-life for the total US economy is equivalent to the value obtained in Papell et al. (2000) for yearly data.

Thus, the conclusion that could be drawn from this …rst analysis is that the unemployment rate is not stationary for an signi…cant number of US States. This implies that we can also accept the unit root null hypothesis for the total US unemployment rate. However, according to the results of previous papers, as well as that emerge from our descriptive analysis, it is possible for these rates to be a¤ected by the presence of some structural breaks. If we want to account for these breaks, it seems to be appropriate to use those statistics that allow for the presence of breaks in the model speci…cation; otherwise, the inference could be biased towards the acceptance of the unit root null hypothesis.

Thus, the results for the case where the model speci…cation includes a single break are reported in Table 2. A simple inspection of this Table leads us to conclude that the evidence against the unit root null hypothesis is clearly greater than in the preceding case. The results contained in this Table provide strong evidence against the unit root hypothesis for a total of 38 States, whilst it is accepted for just 7 States. The inclusion of a single break also implies a signi…cant fall in the half-life measures.

When we analyze the behavior of the total US unemployment rate, we cannot reject the unit root null hypothesis. Nevertheless, we can observe some reduction in the half-life measure when including a single break and, furthermore, the estimation of the period where this break occurs is congruent with the estimations of this parameter for the rest of the US States. In spite of this, the unit root null hypothesis is clearly supported by the data.

If we now consider the possible presence of a second break in the evolution of the unemployment rates, the evidence against the unit root null hypothesis is even greater. According to the results of Table 3, we can now strongly reject this hypothesis for 45 States. Furthermore, the half-life measure is again reduced in value. Thus, we can strongly reject the unit root null hypothesis for all the States under consideration, save for the cases of Alabama, Minnesota, Missouri and Oregon.

As regards the breaks themselves, we can see that the most relevant breaks are those which occurred in the early 1980's, in the early 1990's, and in the mid 1990's. The first reflects the recovering of the US economy as it emerges from the depressive cycle caused by the 2nd oil crisis, whilst the second also indicate the recovering of the US economy in the earlier 1990's. Both periods of change coincide with that used in Partridge and Rickman (1997). The break related to the mid 1990's is clearly related to the acceleration that took placed in the US economy at that time. Again, we can find an economic explanation of this change in Runner (1996), where it is claimed that an extensive modification in the insurance legislation might be the cause of the change in the unemployment rate. When just a single break is considered, we find that most of these are related to the recession caused by the 2nd oil crisis. More precisely, the value of the estimated time of the break belongs to the (1979 : 4, 1982 : 4) interval for 32 States. By contrast, the estimation of this parameter falls within the (1990 : 1, 1992 : 4) interval for 16 States, whilst the estimation is around 1995 for only two States. When we consider a second break, the estimation of the parameter TB_1 belongs to the (1979 : 4, 1982 : 4) interval for 47 States. The estimation of TB_2 is a little more disperse and this estimator belongs to the (1990 : 1, 1992 : 2) interval for 35 States, whilst it lies in the (1995 : 1, 1997 : 4) interval for 13 States. As we can infer from these results, there is a significant amount of dispersion in the estimation of the data where the breaks have occurred, in spite of these breaks being associated to periods of time when it is known that they exercise a clear influence over the evolution of the US economy.

Another important insight is that the magnitudes of the breaks that can be associated to the post 2^{nd} oil crisis period always have a negative sign. Moreover, they are usually greater in absolute value than the rest of the break magnitudes. Thus, the expansionist cycle which followed the 2^{nd}oil crisis involved a significant reduction in the unemployment rates for the US States. Similarly, the 1997 break reduced the unemployment level, albeit to a lesser degree. The conclusion for the 1990 break is not so clear, in that we can appreciate a different influence over the individual States. For example, the unemployment rate in Florida increased after 1990 : 4, whilst in Georgia it fell during exactly the same period.

Finally, we should comment on the results for the total US unemployment rate. Here, we can observe that the presence of the second break leads us to reject the unit root null hypothesis. Similarly, we can also see a significant reduction in the half-life measure. These results are apparently in agreement with those obtained in the analysis of the individual States. However, there are two aspects which cannot be so easily interpreted in the light of the disaggregated results. First, we can see that the half-life measure, in spite of having experienced a significant fall, is markedly higher than the other values reported in Table 4. Thus, the rejection of the unit root hypothesis cannot hide the existence of a great degree of persistence in the evolution of the total US unemployment rate.

Secondly, the estimations of the time of the breaks for the total US unemployment rate are both related to the evolution of the economy in the early 1980's. These breaks are separated by only 5 quarters. This is not habitual in the analysis of the individual States, where it has been frequent to find one break at the beginning of the sample and a second located in the middle or at the end.

Thus, if we had analyzed the evolution of the total US unemployment rate outside of this regional framework, it would not have been an easy task to explain these results. Indeed, we might have considered them to be rather confusing. However, if we take into account the analysis of the individual US States, it becomes much easier to offer a plausible explanation of these apparently odd results. The key to this puzzle is the presence of a significant amount of heterogeneity in the behavior of the unemployment rate corresponding to these States. This is understandable if we bear in mind that the total US unemployment rate can be thought of as a weighted average of the behavior of unemployment in the different States. In our case, we have offered evidence that the behavior of this variable for each State is quite related, but certainly cannot be considered as homogeneous. Thus, for example, we can admit the presence of a break in the early 1990's, but we cannot offer a unique estimation of the period of time where this break appears. Similarly, for the post 2^{nd} oil crisis break, we know that this effect has had an influence over the evolution of the unemployment rate in the US States. However, this influence has appeared at different periods of time and, furthermore, with distinct levels of intensity. The results for the total US unemployment rate are simply reflecting this heterogeneity. Therefore, the appearance of two breaks very close together in time should be understood as the existence of a phenomenon which has altered the evolution of the unemployment rate, but whose effect has impinged on each of the States in different periods of time. Similarly, given that the impact of the breaks on the individual States has not been homogenous, its consequence is that the variable tends to adjust to the inclusion of two breaks in the model speci…cation and, therefore, the estimation exhibits both a positive and a negative break related to the same time period. It might be possible that the inclusion of an additional break could help us to clarify the interpretation of this dual behavior at the beginning of the 1980's.

Thus, these results o¤er further evidence on the problems that aggregation can cause in the study of time properties. It is clear that the most appropriate way to solve these kinds of problem is to carry out a dual analysis. That is to say, whilst it is of interest to analyze the behavior of the total US unemployment rate, it is also advisable to explore the behavior of the components of this ratio. Consequently, larger volumes of disaggregated data are extremely useful; otherwise, and as we have con…rmed, the researcher may be induced to draw quite misleading conclusions.

Have analyzed the time properties of US unemployment, it is now appropriate to carry out some additional studies in order to extract their economic implications. This is the aim of the next Section.

4 Economic implications of the empirical results

As is now clearly established, the main aim of this paper is to analyze the structure of the US unemployment rate, paying particular attention to the time properties of these series for the individual US States. We have therefore focused our e¤ort on determining both the integration order of these rates and on detecting the existence of shifts in their evolution. These results obviously have some economic implications that are worthy of discussion. In particular, we consider it of great interest to obtain the natural rate of unemployment of the individual States in question.

To that end, let us begin by reminding ourselves that the results presented in the above Section have led us to conclude that the unemployment rates of these States do not exhibit a signi…cant amount of persistence. By contrast, we are in a position to strongly reject the presence of a unit root for these variables. This result has some signi…cance from a theoretical point of view, in that it is implicitly assuming that our data base does not support those theoretical models which rest on the concept of complete hysteresis. Rather, there is greater evidence in favor of the so-called structuralist models, where the permanent shift in the long-term unemployment rate are no caused by previous unemployment. Coakley et al. (2001) o¤er an interesting discussion about these two sets of explanation for changes in long-term unemployment.. These authors describe the two schools as the persistence and the structuralist school, respectively. Thus, and given that these latter models assume the presence of an equilibrium unemployment rate, it is appropriate to estimate the value of this equilibrium rate, also known as NAIRU, for the different unemployment rates of those individual US States that offer evidence against the unit root null hypothesis. Here, we adopt a similar procedure to that employed in Papell et al. (2000), which estimates this value for some OECD countries. Let us consider the theoretical model proposed in Layard et al. (1991). This model, included in the structuralist line of thinking, is not based on labor demand and supply functions, but rather on equations for the fixing of prices and wages. In this context, the prices depend on a mark-up on expected wages, whilst the wages depend on a mark-up on expected prices. Consequently, the two functions are:

$$p - w^e = \beta_0 - \beta_1 u \tag{2}$$

$$w - p^e = \gamma_0 - \gamma_1 u \tag{3}$$

where p (prices) and w (nominal wage) are expressed in logs, u is the unemployment rate, β_0 and γ_0 are the mark-up or pulse parameters and β_1 and γ_1 the flexibility parameters. We consider that $\beta_0, \beta_1, \gamma_0, \gamma_1 > 0$, with these parameters being determined by different factors.

Adopting the common assumptions, we consider that both prices and nominal wages follow a random walk. Consequently, in the long term $w^e = w$ and $p^e = p$. From (2) and (3), we can obtain the unemployment equilibrium rate:

$$u^* = \frac{\beta_0 + \gamma_0}{\beta_1 + \gamma_1} \tag{4}$$

Of course, this unemployment rate can be modified due to changes in the structural parameters that are associated to institutional factors, such as changes in the minimum wage, in unemployment insurance, etc. If we admit, as is usually the case, the presence of nominal surprises, then we allow for the existence of some errors in the creation of the expectations and, therefore, $w^e - w = p - p^e \neq 0$. Thus, we can show that:

$$u = u^* - \frac{p - p^e}{\theta_1} \tag{5}$$

where $\theta_1 = \frac{\beta_1 + \gamma_1}{2}$. Consequently, the short-term unemployment rate movements depend on the existence of an acceleration or deceleration process in prices or, equivalently, in nominal wages. If we suppose that the inflation follows a random walk and, therefore, that the variations in inflation are stationary, we have that:

$$\Delta p = \Delta p_{-1} + \varepsilon \tag{6}$$

where ε is a stationary noise. We should note that this is not a very restrictive assumption, given that Laubach (2000) and Roberts and Morin (1999) offer empirical support for this result in the case of the USA. Then, it is true that:

$$p^e = E(p) = p_{-1} - \Delta p_{-1} \tag{7}$$

and we can obtain the classical Phillips curve, which can be stated as follows:

$$u = u^* - \frac{\Delta p - \Delta p_{-1}}{\theta_1} \tag{8}$$

Thus, we can observe that the movements of the unemployment rate around the natural rate depend on the evolution of prices or, given our initial assumptions, of nominal wages. The last equation can be alternatively written as follows:

$$\Delta p = \Delta p_{-1} - \theta_1 (u - u^*) \tag{9}$$

proving that the inflation depends on the deviation of the unemployment rate from its natural rate. However, and as is pointed out in Layard et al. (1991), empirical application shows that the history of unemployment is important in explaining the behavior of inflation. Consequently, it is advisable to reformulate (9) by considering the dynamics effect of unemployment. This leads us to the following relationship:

$$\Delta p = \Delta p_{-1} - \theta_1 (u - u^*) - \theta_{11} (u - u_{-1}) \tag{10}$$

which can be expressed in terms of u as follows:

$$u = \frac{\theta_1}{\theta_1 + \theta_{11}} u^* + \frac{\theta_{11}}{\theta_1 + \theta_{11}} u_{-1} - \frac{1}{\theta_1 + \theta_{11}} (\Delta p - \Delta p_{-1}) \tag{11}$$

where we should consider that $\frac{\theta_{11}}{\theta_1+\theta_{11}}$ measures the persistence of the unemployment rate. Moreover, as we have assumed inflation to be stationary, if $\left|\frac{\theta_{11}}{\theta_1+\theta_{11}}\right| < 1$, the unemployment rate will also be stationary and, consequently, it is possible to determine the value of the equilibrium unemployment rate.

In our case, this is a simple task, in that we only need to calculate the value of u^*. We can do this by simply comparing (11) with (1), observing that we can identify various elements. For example, we can see directly that $\frac{\theta_{11}}{\theta_1+\theta_{11}}$ corresponds to ρ, which implicitly implies that $\frac{\theta_1}{\theta_1+\theta_{11}} = 1 - \rho$. Similarly, it is also true that $\frac{\theta_1}{\theta_1+\theta_{11}} u^*$ is related to $\alpha + \sum_{i=1}^{n} d_i DU_{it}$. Therefore, the NAIRU can be stated as follows:

$$u^* = \frac{\theta_1 + \theta_{11}}{\theta_1}\left(\alpha + \sum_{i=1}^{n} d_i DU_{it}\right) = \frac{\alpha + \sum_{i=1}^{n} d_i DU_{it}}{1 - \rho} \qquad (12)$$

This approach allows us to obtain a simple, but accurate, estimation of the long-term NAIRU by only bearing in mind the estimation of the parameters of model (1). Of course, this method does not allow us to analyze the structural determinant of this natural rate. Nevertheless, we should recall that it would be possible to determine the short-term NAIRU from (11) by simply removing the last element of this equation. We have preferred to focus exclusively on the long-term NAIRU given that this can provide us with much more interesting insights than the short-term NAIRU, at least in the framework adopted in this paper.

The estimation of the NAIRU for the 46 States which rejected the unit root null hypothesis is reported in Table 5. We have considered $n = 2$, in that we have found strong evidence in favor of the existence of a double break in the evolution of the unemployment rate. Therefore, there are three different segments in the evolution of the NAIRU for these States, although we can exclusively use the last of these for the purposes of comparison, given that the time breaks are not coincident.

A first examination of the results suggests that there has been a generalized and significant reduction in the NAIRU of the States. However, this reduction has not been linear; rather, it has been produced in an heterogenous manner. The most noteworthy example is that of Michigan, which shows a NAIRU of 8.61% at the beginning of the sample and one of 4.29% at the end. This implies that the unemployment rate of this particular State has experienced a reduction of more than 4 points, implying 50% of the original NAIRU value. Moreover, it has the lowest NAIRU at the end of the sample. The States of Maryland and West Virginia also show similar reductions. By contrast, the estimation of the NAIRU for other States, such as Maine or New Hampshire, has not demonstrated a significant variation. In these two cases, the NAIRU has decreased by around 1 point, that is to say, some 20% of their respective NAIRU at the beginning of the sample.

We should also note that the estimations of the NAIRU at the end of the sample lie on the (4.32%; 5.96%) interval. As we have just mentioned, Michigan reports the minimum value, followed by California (4.42%) and Minnesota (4.46%). It is also worthy of note that the States located on the

Paci…c Coast (California and Washington) show a relatively low estimation of the NAIRU. So far as the States with a higher estimation of the NAIRU are concerned, we should …rst observe that Louisiana presents the highest value, followed by Florida (5:94%), Utah (5:90%) and Kansas (5:86%). Thus, macroeconomic policy targeted towards in‡uencing the US labor, market may have varying e¤ects on the natural unemployment rates of the States.

Finally, we should o¤er some information on the value of the NAIRU for the total USA. However, this is not a simple task in that, given the results of the previous section, it does not seem to be advisable to obtain this value by following the same procedure as employed in the case of the individual States. The value obtained following such a procedure would be 5:43%, but we should nevertheless recall that we have considered the ADF speci…cation with $n = 2$ as being insu¢ cient to capture the behavior of the unemployment rate.

Another alternative would be to use the mean of the estimations of the NAIRU at the end of the sample. In this case, the NAIRU value for the total USA would be 5:31%. The problem here is that the weight of each State is not same, and thus this value could be an ine¢ cient approximation. However, and given that the standard deviation of the estimations of the NAIRU is quite low (0:42), we consider that the values obtained under the two alternatives o¤er a good approximation to the true value of the total US NAIRU. Note that both values are almost identical and quantitatively similar to that obtained in Papell et al. (2000) and Laubach (2000).

5 Conclusions

This paper has analyzed the time properties of the unemployment rates for both the individual US States and for the USA as a whole. To that end, we have used those unit root tests which allow for the presence of some breaks in the evolution of the variable being studied. The inclusion of these breaks has been shown to be determinant in describing the structure of the US unemployment rate.

On this basis, we have found that we can reject the unit root null hypothesis for a total of 46 States whenever two breaks are included in the model speci…cation. Although the estimations of the times when these breaks occur are meaningful from an economic point of view, we cannot consider them as showing a homogenous pattern of behavior. By contrast, both the estimation of the times of the breaks and their magnitude are relatively heterogenous, providing evidence that economic phenomena do not a¤ect all the States in an equivalent manner.

The importance of this first result is better understood when we analyze the total US unemployment rate. The results for this variable are not entirely conclusive: on the one hand, we can reject the unit root null hypothesis when two breaks are added to the model specification; on the other, the inclusion of these breaks does not avoid the presence of a high degree of persistence. Furthermore, the estimations and the magnitudes of the breaks are not congruent with the results obtained for the individual States. Thus, we can interpret these apparently opposing results from the point of view of a aggregation problem. On this basis, our advice is to carry out a dual analysis in this type of scenario, that is to say, an analysis for the total US unemployment plus an that of the individual States. Otherwise, we run the risk that the conclusions could be quite distorted.

Finally, given that the rejection of the unit root null hypothesis can be considered as evidence in favor of the structuralist labor model, we have offered a simple estimation of the NAIRU for each of the US States. The results are quite homogeneous and the estimation of this natural rate lies on the $(4{:}32\%; 5{:}96\%)$ interval. Using the sample mean of this regional NAIRU as an approximation to the total US NAIRU, this value (5:31%) is qualitatively similar to that reported in recent papers.

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6 Appendix A. Tables

Table 1. Testing for unit roots in the US employment rate. No break case

States	$\rho - 1$	τ_1	ϕ	HL
Alabama	-0.42	-2.44	2.55	1.27
Alaska	-0.68	-6.81^a	4.42	0.61
Arizona	-0.43	-2.39	2.51	1.23
Arkansas	-0.58	-5.86^a	3.74	0.80
California	-0.41	-2.36	2.43	1.31
Colorado	-0.62	-6.20^a	4.01	0.72
Connecticut	-0.50	-3.03^c	3.00	1.00
Delaware	-0.67	-6.52^a	4.22	0.63
District Columbia	-0.48	-3.01^c	2.94	1.06
Florida	-0.53	-4.24^a	3.42	0.92
Georgia	-0.54	-4.13^a	3.24	0.89
Idaho	-0.49	-3.96^a	3.20	1.03
Illinois	-0.46	-2.81^c	2.79	1.12
Indiana	-0.53	-4.37^a	3.51	0.92
Iowa	-0.84	-7.82^a	5.07	0.38
Kansas	-0.69	-6.71^a	4.52	0.59
Kentucky	-0.43	-2.20	2.56	1.23
Louisiana	-0.63	-6.26^a	4.13	0.70
Maine	-0.52	-3.03^c	3.13	0.94
Maryland	-0.48	-3.97^a	3.09	1.06
Massachusetts	-0.44	-2.72^c	2.59	1.20
Michigan	-0.43	-3.70^b	2.74	1.23
Minnesota	-0.38	-2.53	2.26	1.45
Mississippi	-0.54	-4.35^a	3.41	0.89
Missouri	-0.51	-3.21^c	3.08	0.97
Montana	-0.72	-7.12^a	4.65	0.54
Nebraska	-0.44	-2.39	2.61	1.20
Nevada	-0.59	-5.93^a	3.81	0.78
New Hampshire	-0.41	-2.30	2.37	1.31
New Jersey	-0.61	-6.12^a	3.94	0.74
New Mexico	-0.39	-2.26	2.27	1.40
New York	-0.65	-6.37^a	4.12	0.66
North Carolina	-0.49	-3.04^c	2.97	1.03
North Dakota	-0.73	-6.96^a	4.62	0.53
Ohio	-0.40	-2.65	2.41	1.36
Oklahoma	-0.53	-4.19^a	3.42	0.92

States	$\hat{\rho}-1$	τ_{μ}	$\hat{\mu}$	HL
Oregon	-0.35	-2.46	2.08	1.61
Pennsylvania	-0.48	-3.99^{a}	3.12	1.06
Rhode Island	-0.44	-2.82^{c}	2.65	1.20
South Carolina	-0.67	-6.60^{a}	4.34	0.63
South Dakota	-0.41	-2.20	2.45	1.31
Tennessee	-0.66	-6.43^{a}	4.27	0.64
Texas	-0.40	-2.08	2.35	1.36
Utah	-0.50	-4.12^{a}	3.22	1.00
Vermont	-0.45	-2.82^{c}	2.67	1.16
Virginia	-0.46	-3.85^{a}	2.93	1.12
Washington	-0.44	-2.77^{c}	2.63	1.20
West Virginia	-0.52	-4.25^{a}	3.28	0.94
Wisconsin	-0.40	-2.69	2.42	1.36
Wyoming	-0.58	-4.46^{a}	3.68	0.80
USA	-0.04	-1.67	0.23	16.98

Table 1(cont.). Testing for unit roots in the US employment rate. No break case

This Table includes the ADF statistic used for testing for the presence of a unit root in the unemployment rate of the US States reported in the first column. The value of the truncation lag parameter was chosen according to the k(t) method proposed in Ng and Perron (1995), using kmax = 5. We have tabulated the critical values of this case. To that end, we generated 50,000 replications of a white noise and, subsequently, have estimated the model (1) with $n = 0$. These critical values are -3.76, -3.06 and -2.71 for a 1%, 5%and 10% significance level, respectively. HL stands for the half-life measure of persistence, which is obtained as $\ln 0.5/\ln\hat{\rho}$

[c], [b], [a] means 10%, 5% and 1% rejection of the null hypothesis, respectively.

Table 2. Testing for unit roots in the US employment rate. n=1

States	$\hat\rho - 1$	$\max \Phi_1$	$\hat\lambda$	$\hat d_1$	TB_1	HL
Alabama	-0.74	8.28	5.67	-1.38	82.2	0.51
Alaska	-0.81	28.76[a]	6.92	-2.00	81.1	0.42
Arizona	-0.75	25.32[a]	4.79	-0.98	92.2	0.50
Arkansas	-0.62	23.11[a]	4.62	-0.76	80.1	0.72
California	-0.62	9.08	4.10	-1.01	90.4	0.72
Colorado	-0.71	23.64[a]	6.06	-1.76	80.3	0.56
Connecticut	-0.94	11.53[b]	7.74	-2.47	82.1	0.25
Delaware	-0.67	25.51[a]	4.42	-0.77	91.4	0.63
District Columbia	-0.83	30.06[a]	6.25	-1.53	81.4	0.39
Florida	-0.69	26.34[a]	4.58	-0.75	91.3	0.59
Georgia	-0.63	11.61[b]	4.80	-1.20	80.3	0.70
Idaho	-0.57	12.78[b]	3.87	-0.77	91.3	0.82
Illinois	-0.66	7.25	5.52	-1.79	81.2	0.64
Indiana	-0.69	32.80[a]	4.67	-0.75	91.1	0.59
Iowa	-0.90	35.44[a]	6.66	-1.48	81.1	0.30
Kansas	-0.72	39.38[a]	4.91	-0.72	90.4	0.54
Kentucky	-0.63	6.98	6.1	-2.57	80.4	0.70
Louisiana	-0.67	26.71[a]	4.53	-0.59	80.3	0.63
Maine	-0.82	28.81[a]	5.98	-1.26	82.2	0.40
Maryland	-0.82	30.03[a]	6.93	-2.05	82.1	0.40
Massachusetts	-0.70	12.68[b]	5.49	-1.54	81.3	0.58
Michigan	-0.61	11.15[b]	5.45	-1.86	81.3	0.74
Minnesota	-0.59	10.23[c]	3.75	-1.15	95.4	0.78
Mississippi	-0.85	31.20[a]	7.23	-2.11	81.2	0.37
Missouri	-0.61	8.27	3.94	-0.72	91.1	0.74
Montana	-0.84	31.28[a]	7.09	-1.92	81.1	0.38
Nebraska	-0.83	29.60[a]	5.26	-1.47	95.3	0.39
Nevada	-0.70	23.15[a]	6.43	-2.17	80.4	0.58
New Hampshire	-0.82	14.75[a]	6.06	-1.51	82.2	0.40
New Jersey	-0.65	24.78[a]	4.78	-0.79	79.4	0.66
New Mexico	-0.71	8.60	5.67	-1.76	82.1	0.56
New York	-0.66	24.51[a]	4.46	-0.86	91.4	0.64
North Carolina	-0.80	27.98[a]	5.99	-1.49	81.4	0.43
North Dakota	-0.73	29.48[a]	4.82	-0.82	91.3	0.53
Ohio	-0.62	10.71[c]	4.69	-1.15	82.4	0.72
Oklahoma	-0.71	28.54[a]	4.79	-0.85	91.2	0.56

Table 2(cont). Testing for unit roots in the US employment rate. n=1						
States	$\hat{\rho}-1$	$\max \Phi_1$	$\hat{\mu}$	$\hat{d}_1$	TB_1	Half-life
Oregon	-0.51	8.25	3.28	-0.70	92.1	0.97
Pennsylvania	-0.55	14.23^{a}	3.84	-0.86	91.1	0.87
Rhode Island	-0.82	31.25^{a}	6.25	-1.59	81.1	0.40
South Carolina	-0.71	37.50^{a}	4.83	-0.81	90.4	0.56
South Dakota	-0.82	29.45^{a}	5.26	-0.99	92.2	0.40
Tennessee	-0.70	32.49^{a}	4.72	-0.69	90.3	0.58
Texas	-0.78	27.63^{a}	5.62	-1.20	81.3	0.46
Utah	-0.77	28.10^{a}	6.33	-1.74	81.3	0.47
Vermont	-0.63	10.20^{c}	5.87	-2.31	80.2	0.70
Virginia	-0.79	27.62^{a}	6.79	-2.12	81.4	0.44
Washington	-0.63	10.55^{c}	4.76	-1.15	81.1	0.70
West Virginia	-0.80	28.01^{a}	6.96	-2.16	81.2	0.43
Wisconsin	-0.57	10.19^{c}	4.41	-1.12	80.4	0.82
Wyoming	-0.87	33.07^{a}	7.21	-1.91	81.1	0.34
USA	-0.05	8.91	0.50	-0.27	82.3	13.51

This Table reports the results obtained from the estimation of model (1) when n=1. The columns of this Table reflect, for each State, the value of the estimation of the parameter $\alpha = \hat{\rho} - 1$, the value of the $\max \Phi_1$ statistic, the estimation of the intercept of the model ($\hat{\mu}$), the estimation of the dummy which measures the impact of the break ($\hat{d}_1$), the estimation of the period where this break occurs (TB_1) and the half-life measure of persistence (HL), which is obtained as $\ln 0.5 / \ln \hat{\rho}$. The value of the lag truncation parameter was chosen according to the k(t) method, proposed in Ng and Perron (1995), using kmax = 5. The critical values for $\max \Phi_1$ are 9.72, 11.05 and 13.90 respectively, for the 10%, 5% and 1% level of significance.

[c], [b], [a] means 10%, 5% and 1% rejection of the null hypothesis, respectively.

Table 3. Testing for unit roots in the US unemployment rate. n=2

States	½ ¡ 1	max ©$_2$	^	$\hat{d}_1$	TB_1	$\hat{d}_2$	TB_2	HL
Alabama	¡ 0:82	9:94	6:67	¡ 1:77	80:4	¡ 0:47	91:1	0:40
Alaska	¡ 0:83	21:76[a]	7:14	¡ 1:94	81:1	¡ 0:88	95:3	0:39
Arizona	¡ 0:90	22:64[a]	6:55	¡ 1:16	80:3	¡ 0:73	91:1	0:30
Arkansas	¡ 0:64	17:57[a]	4:74	¡ 0:60	80:1	¡ 0:66	92:2	0:68
California	¡ 0:96	12:29[b]	7:18	¡ 1:56	82:2	¡ 1:38	97:2	0:22
Colorado	¡ 0:65	19:45[a]	4:50	¡ 0:27	79:4	¡ 0:65	92:1	0:66
Connecticut	¡ 0:92	9:97	6:70	¡ 1:23	83:2	¡ 0:82	92:4	0:27
Delaware	¡ 0:67	19:90[a]	4:86	¡ 0:80	82:1	¡ 0:41	91:4	0:63
District Columbia	¡ 0:88	23:00[a]	6:65	¡ 1:72	81:4	0:03	90:2	0:33
Florida	¡ 0:84	20:63[a]	6:41	¡ 1:54	82:4	0:12	90:4	0:38
Georgia	¡ 0:85	21:32[a]	6:51	¡ 1:68	81:3	¡ 0:19	90:4	0:37
Idaho	¡ 0:77	21:85[a]	5:93	¡ 1:23	83:3	¡ 0:38	91:2	0:47
Illinois	¡ 0:87	23:31[a]	6:51	¡ 1:45	81:2	¡ 0:28	91:1	0:34
Indiana	¡ 0:78	27:14[a]	6:21	¡ 1:55	82:2	¡ 0:22	91:1	0:46
Iowa	¡ 0:92	26:93[a]	6:80	¡ 1:53	81:1	¡ 0:12	90:4	0:27
Kansas	¡ 0:83	36:51[a]	6:86	¡ 1:94	82:1	¡ 0:06	90:4	0:39
Kentucky	¡ 0:89	22:56[a]	6:42	¡ 1:00	82:3	¡ 0:81	92:2	0:31
Louisiana	¡ 0:78	24:62[a]	6:36	¡ 1:66	81:4	¡ 0:05	90:3	0:46
Maine	¡ 0:81	21:60[a]	5:48	¡ 0:31	80:1	¡ 0:59	85:1	0:42
Maryland	¡ 0:87	22:19[a]	7:39	¡ 1:99	82:1	¡ 1:44	97:2	0:34
Massachusetts	¡ 0:89	23:01[a]	6:59	¡ 1:35	81:4	¡ 0:87	95:1	0:31
Michigan	¡ 0:85	20:43[a]	7:32	¡ 2:08	81:4	¡ 1:57	97:1	0:37
Minnesota	¡ 0:65	9:02	4:76	¡ 0:78	80:1	¡ 1:08	95:4	0:66
Mississippi	¡ 0:88	24:45[a]	7:52	¡ 1:99	81:2	¡ 1:25	95:4	0:33
Missouri	¡ 0:68	8:93	5:09	¡ 0:95	80:4	¡ 0:63	91:2	0:61
Montana	¡ 0:80	25:52[a]	5:71	¡ 0:40	80:2	¡ 0:99	92:3	0:43
Nebraska	¡ 0:91	30:01[a]	6:79	¡ 1:38	80:3	¡ 0:51	90:4	0:29
Nevada	¡ 0:65	19:33[a]	4:87	¡ 0:64	80:1	¡ 0:73	92:2	0:66
New Hampshire	¡ 0:77	17:69[a]	5:00	0:04	88:1	¡ 1:11	92:4	0:47
New Jersey	¡ 0:67	19:25[a]	4:91	¡ 0:60	79:4	¡ 0:71	92:1	0:63
New Mexico	¡ 0:98	13:28[b]	7:55	¡ 1:82	82:1	¡ 1:30	96:1	0:18
New York	¡ 0:68	18:79[a]	5:09	¡ 0:94	82:1	¡ 0:47	91:4	0:61
North Caroline	¡ 0:86	20:91[a]	6:53	¡ 1:47	81:4	¡ 0:42	90:3	0:35
North Dakota	¡ 0:72	22:40[a]	5:18	¡ 0:77	81:4	¡ 0:46	91:3	0:54
Ohio	¡ 0:83	20:60[a]	6:49	¡ 1:56	81:3	¡ 1:21	96:2	0:39
Oklahoma	¡ 0:76	21:99[a]	6:02	¡ 1:35	82:3	¡ 0:39	91:2	0:49

Table 3(cont.). Testing for unit root in the US employment rate. n=2								
States	$\rho - 1$	$\max \Phi_2$	μ	$\hat{d}_1$	TB_1	$\hat{d}_2$	TB_2	HL
Oregon	-0.69	10.12	5.61	-1.63	81.2	-0.42	91.1	0.59
Pennsylvania	-0.77	24.22[a]	6.33	-1.76	82.4	-0.31	91.1	0.47
Rhode Island	-0.85	24.14[a]	6.46	-1.59	81.1	-0.26	90.4	0.37
South Caroline	-0.83	35.73[a]	6.91	-2.05	82.1	-0.13	90.4	0.39
South Dakota	-0.83	22.93[a]	5.93	-1.21	81.4	-0.15	90.3	0.39
Tennessee	-0.82	30.00[a]	6.72	-1.86	81.4	-0.10	90.3	0.40
Texas	-0.85	20.57[a]	6.18	-1.06	82.2	-0.71	92.1	0.37
Utah	-0.81	22.41[a]	6.81	-2.19	82.1	0.16	90.2	0.42
Vermont	-0.80	19.50[a]	5.95	-1.21	82.1	-0.79	95.1	0.43
Virginia	-0.83	21.37[a]	7.14	-2.27	81.4	-0.04	90.3	0.39
Washington	-0.86	21.62[a]	6.50	-1.36	81.3	-1.05	95.4	0.35
West Virginia	-0.82	20.91[a]	7.17	-2.05	81.2	-1.11	95.4	0.40
Wisconsin	-0.75	16.15[a]	5.50	-0.94	82.1	-0.57	93.4	0.50
Wyoming	-0.83	26.63[a]	5.94	-0.62	80.2	-1.21	95.3	0.39
USA	-0.07	12.06[b]	0.53	0.47	81.2	-0.62	82.3	9.55

This Table reports the results obtained from the estimation of model (1) when n=2. The columns of this Table reflect, for each State, the value of the estimation of the parameter $\alpha = \rho - 1$, the value of the $\max \Phi_2$ statistic, the estimation of the intercept of the model (μ), the estimation of the dummies which measure the impact of the break ($\hat{d}_i$, $i = 1, 2$), the estimation of the periods where these breaks occur (TB_i, $i = 1, 2$) and the half-life measure of persistence (HL), which is obtained as $\ln 0.5 / \ln \rho$. The value of the lag truncation parameter was chosen according to the k(t) method, proposed in Ng and Perron (1995), using kmax = 5. The critical values for $\max \Phi_2$ are 10.61, 11.64 and 13.88, respectively, for the 10%, 5% and 1% levels of significance.

[c], [b], [a] means 10%, 5% and 1% rejection of the null hypothesis, respectively.

Table 4. Evolution of the NAIRU for the US States

State	$u_1^{¤}$	$u_2^{¤}$	$u_3^{¤}$
Alabama	8:13	5:98	5:40
Alaska	8:60	6:27	5:20
Arizona	7:28	5:99	5:18
Arkansas	7:41	6:47	5:44
California	7:48	5:85	4:42
Colorado	6:92	6:51	5:51
Connecticut	7:28	5:95	5:05
Delaware	7:25	6:06	5:45
District Columbia	7:56	5:60	5:64
Florida	7:63	5:80	5:94
Georgia	7:66	5:68	5:46
Idaho	7:70	6:1	5:61
Illinois	7:48	5:82	5:49
Indiana	7:96	5:97	5:69
Iowa	7:39	5:73	5:60
Kansas	8:27	5:93	5:86
Kentucky	7:21	6:09	5:18
Louisiana	8:15	6:03	5:96
Maine	6:77	6:38	5:65
Maryland	8:49	6:21	4:55
Massachusetts	7:40	5:89	4:91
Michigan	8:61	6:16	4:32
Minnesota	7:32	6:12	4:46
Mississippi	8:55	6:28	4:86
Missouri	7:49	6:09	5:16
Montana	7:14	6:64	5:40
Nebraska	7:46	5:95	5:38
Nevada	7:49	6:51	5:38
New Hampshire	6:49	6:55	5:10
New Jersey	7:33	6:43	5:37
New Mexico	7:70	5:85	4:52
New York	7:49	6:10	5:41
North Carolina	7:59	5:88	5:40
North Dakota	7:19	6:13	5:49
Ohio	7:82	5:94	4:48
Oklahoma	7:92	6:14	5:63

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Table 4 (cont.). Evolution of the NAIRU for the US States.			
State	u_1^*	u_2^*	u_3^*
Oregon	8:13	5:77	5:16
Pennsylvania	8:22	5:94	5:53
Rhode Island	7:60	5:73	5:42
South Carolina	8:33	5:86	5:70
South Dakota	7:14	5:69	5:51
Tennessee	8:20	5:93	5:80
Texas	7:27	6:02	5:19
Utah	8:41	5:70	5:90
Vermont	7:44	5:93	4:94
Virginia	8:60	5:87	5:82
Washington	7:56	5:98	4:76
West Virginia	8:74	6:24	4:89
Wisconsin	7:33	6:08	5:32
Wyoming	7:16	6:41	4:95

This Table reports the values of the NAIRU (in percentages) for the 46 US States which have led us to reject the unit root null hypothesis. These values have been calculated from the estimation of model (1) with n = 2. The periods where the NAIRU changes are reported in Table 3 and 4 for the n = 1 and n = 2 cases, respectively.