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What's 'structural' about unemployment in Europe: On the Determinants of the European Commission's NAIRU Estimates

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What's Structural About Unemployment in Europe? On the Determinants of the European Commission's NAIRU Estimates

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Abstract

This paper analyzes the determinants of the European Commission's NAIRU estimates for 14 European OECD countries during 1985-2012. The NAIRU is a poor proxy for 'structural unemployment': Labor market institutions – employment protection legislation, union density, tax wedge, minimum wages – underperform in explaining the NAIRU, while cyclical variables – capital accumulation and boom-bust patterns in housing markets – play an important role. This is relevant since the NAIRU is used to compute potential output and structural budget balances and, hence, has a direct impact on the scope and evaluation of fiscal policy in EU countries.

JEL codes: C54, E24, E62
Keywords: NAIRU, potential output, fiscal policy coordination, unemployment in Europe, labor market flexibility

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1 Introduction

The non-accelerating inflation rate of unemployment – or, in short, NAIRU – is a major concept in modern macroeconomics. Its core proposition is that, for any economy and at any point in time, there exists some (unobserved) rate of unemployment at which inflation remains constant. Historically, the NAIRU can be seen as a direct offspring of the famous Phillips curve, which posits a negative relationship between unemployment and (changes in) inflation. However, over time the NAIRU has also been identified with the idea of a natural rate of unemployment (e.g. Ball and Mankiw, 2002), which would prevail in the absence of any cyclical fluctuations and, hence, represents structural unemployment existing independently of all temporary and seasonal fluctuations (Friedman, 1968; Phelps, 1967). By encapsulating two important concepts of 20th century macroeconomics - the Philips curve as well as the idea of a natural or structural rate of unemployment - the NAIRU has not only been given a great variety of more specific theoretical and econometric interpretations, but has also secured its place as a standard tool in current macroeconomics.

Notwithstanding its success, the NAIRU has also confronted empirical researchers with a troubling question, namely: How to produce reliable empirical estimates of a theoretically postulated but unobservable variable? In many of the past and current applications, this question is resolved by pragmatic approaches, which treat the NAIRU as an unobservable stochastic variable (e.g. Staiger *et al.*, 1997; Franz, 2005; Watson, 2014), employing a variety of econometric models and statistical techniques to estimate this variable. In this paper, we argue that this practice creates a certain tension between theory and empirical application: While theoretical accounts connected with the idea of the 'natural rate of unemployment' posit that structural factors determine the NAIRU, most actual empirical estimations of the NAIRU are devoid of such considerations, but take a comparably empiricist approach, which is either based on pure statistical technique – such as the Hodrick-Prescott filter (Hodrick and Prescott, 1997) – or relies on the integration of a Phillips curve relationship into a statistical de-trending and filtering process, as in typical Kalman Filter applications (e.g. Laubach, 2001; Durbin and Koopman, 2012). These methods are used to separate trend and cyclical components of unemployment without making any reference to the structural factors underlying 'trend unemployment', which is nonetheless interpreted as a suitable estimate for the NAIRU of the economies under study.

From this perspective, the connection between theoretical account and empirical practice is established only implicitly – by effectively assuming that one's de-trended series does indeed represent the structural factors driving unemployment and, hence, is a good proxy of the true NAIRU values in the economy under study. In this paper, we aim to constructively exploit this tension between theory and empirical application by critically assessing the empirical plausibility of the essential underlying hypothesis that the evolution of the NAIRU is driven by structural factors. Specifically, we study whether theoretical arguments on structural unemployment are suitable to empirically interpret commonly used estimates of the NAIRU as an unobservable stochastic variable. In doing so, we assess the robustness and plausibility of these NAIRU estimates and the underlying assumption that these estimates indeed represent the unobservables posited by theory, which is a contribution to answering the question what commonly used NAIRU estimates actually tell us about the state of an economy.

In operationalizing this aim, we focus on a specific case, namely the non-accelerating wage inflation rate of unemployment (NAWRU) as estimated by the European Commission (EC). This case is of major interest for various reasons: First, it employs a Kalman filtering technique based on a Phillips curve framework and, hence, represents a more nuanced case in which the EC tries to go beyond a purely statistical approach and explicitly claims to incorporate a "preference for an economic, as opposed to a statistical, approach" (Havik *et al.*, 2014, p. 5). Second, this case provides a perfect fit with our research question, as the NAWRU is treated as an unobservable stochastic variable. The resulting estimates exist for all EU countries as well as for the USA; hence, they provide an ideal opportunity for systemically analyzing whether there exists a tension between empirical NAWRU estimates and the underlying natural rate theory, which points to the importance of structural labor market features. Third, the EC's NAWRU approach carries exceptionally high policy relevance as the NAWRU is used as a proxy for structural unemployment in calculating cyclically-adjusted budget balances (Havik *et al.*, 2014), which are especially crucial for the coordination of fiscal policy across euro area member states and for the determination of fiscal adjustment paths (e.g. Ecfin, 2013). With high 'structural unemployment', the 'structural component' of the fiscal deficit is estimated to be large. Hence, high NAWRU estimates increase the pressure on EU countries to implement fiscal consolidation measures, because essential fiscal targets in the EU's fiscal regulation framework are set in terms of the structural budget balance. A fourth and final reason why studying the determinants of the NAWRU is highly relevant is that the EC's official estimates are also downloadable from the AMECO database, which is widely used by economic researchers. As a consequence, the theoretical plausibility and robustness of these NAWRU estimates should be of interest to a broader audience, including all those researchers who make use of these data to conduct their own research. Although it might seem to be a rather obvious strategy to econometrically compare actual NAIRU estimates with their supposed theoretical determinants, most current and past research focuses on the empirics of actually observed unemployment (e.g. Blanchard, 2006; Stockhammer and Klär, 2011). According to our best knowledge, there have been only three attempts so far to look at the empirical determinants of Kalman-filtered NAIRU estimates in a larger group of EU countries, where two of these studies were conducted by EC economists themselves (Orlandi, 2012; European Commission, 2013) and one by researchers at the OECD (Gianella *et al.*, 2008). However, our econometric analysis goes beyond this literature in various respects by including additional control variables, by considering a longer time frame – we also include a couple of years after the financial crisis of 2008/2009 – and by providing several additional robustness checks.

Our main empirical finding is that the NAIRU, as estimated by the EC, is a poor proxy for 'structural unemployment'. In our data set of 14 European OECD countries over the time period 1985-2012, institutional labor market indicators – employment protection legislation, union density, minimum wage, tax wedge – do not perform well when it comes to explaining NAIRU estimates. Active labor market policies and unemployment benefit replacement rates are the only statistically significant institutional variables, which in most of our models have the sign expected according to standard theory. We also find that cyclical variables such as capital accumulation and a proxy for boom-bust patterns in housing markets are statistically significant determinants. This finding contradicts the EC's theoretical framework, in which the NAIRU is modeled as the trend component of the unemployment rate, uninfluenced by non-structural factors. We discuss the policy relevance of the econometric findings against the background of the EU's fiscal regulation framework, in which NAIRU estimates are shown to have a direct impact on the scope and evaluation of fiscal policy.

The remainder of this article is structured as follows: In section 2, we provide a short introduction to the empirical estimation and political application of the EC's NAIRU estimates. Section 3 in turn reviews past empirical literature that analyzes the determinants of European unemployment and concisely summarizes the theoretical underpinnings of these applications. In section 4, we develop our basic econometric strategy for assessing the theoretical plausibility and robustness of the EC's NAIRU estimates. Section 5 presents the econometric baseline results and Section 6 assesses the robustness of these findings. Section 7 discusses the role of the NAIRU in theory, empirics and policy practice. Finally, Section 8 concludes our argument.

2 The European Commission's NAIRU approach: Estimation and Application

In accordance with common practice, the EC defines the NAIRU as the unemployment rate at which (wage) inflation remains stable (European Commission, 2014) and, hence, introduces the NAWRU as an alternative acronym for the NAIRU concept[1], which is identified as a proxy for structural unemployment. Moreover, when actual unemployment (u_t) is equal to the NAIRU (N_t) - i.e. the unemployment gap (u_t - N_t) is zero -, the economy is running at potential output (Havik *et al.*, 2014).

The EC's NAIRU model is based on a Kalman Filter applied to an econometric model cast into a state-space framework (Durbin and Koopman, 2012), which consists of (a) a set of assumptions about the unobservables in the model that are of statistical nature (like lag-structures and autoregressive processes), as well as (b) a theoretical component based on a Phillips curve framework. In the latter case, estimated unemployment gaps are used to explain the growth in unit labor costs within the state-space model, possibly in conjunction with a series of exogenous regressors to increase the statistical precision of the underlying Kalman Filter model (Planas and Rossi, 2015). Hence, the theoretical arguments enter the model setup only indirectly to provide additional information for judging between the plausibility of the data and the plausibility of the underlying model within the recursions which make up the Kalman Filter (Kalman, 1960; Harvey, 1990).

[1]In the rest of this paper, the terms NAIRU and NAWRU are, therefore, used interchangeably.

The two so called measurement equations of the NAIRU model formally look as follows:

$$u_t = N_t + G_t \tag{1}$$

$$grulc_t = \alpha grulc_{t-1} + \beta_1 G_t + \beta_2 G_{t-1} + \gamma Z_t + a_{rulct} \tag{2}$$

with $\beta_1 < 0$, $\beta_2 > 0$; and where u_t is the actual unemployment rate; N_t is the trend component of the unemployment rate (i.e., the NAIRU); G_t is the unemployment gap $(u_t - N_t)$; $grulc_t$ is the growth rate of real unit labor costs at time t, and $grulc_{t-1}$ is the lagged growth rate of rulc; Z_t is a vector consisting of exogenous variables (which may include changes in terms of trade and in labor productivity etc.); and a_{rulct} is the error term, which captures measurement errors in $grulc_t$.

Since the Spring Forecast 2014, the EC has been using this Phillips curve specification labeled 'New Keynesian' for most European countries,[2] which is "based on rational expectations [...] [, implying] that a positive unemployment gap [...] is associated with a fall in the growth rate of real unit labor cost" (European Commission, 2014, p. 22). The measurement equations are complemented by a set of state equations, which specify the dynamics of the unobserved components of the model (Planas and Rossi, 2015) and have the following form:

$$\Delta N_t = \eta_{t-1} \tag{3}$$

$$\Delta \eta_t = a_{\eta t} \tag{4}$$

$$G_t = \phi_1 G_{t-1} + \phi_2 G_{t-2} + a_{Gt} \tag{5}$$

where the change in the NAIRU (ΔN_t) is modelled as a Gaussian noise process (η_t) governed by $a_{\eta t}$. All shocks are normally distributed white noises, which are also assumed to be independent from each other. From equations (3) and (4) on the dynamics of the unobserved components, we can see that the NAIRU is specifically modeled as a second-order random walk. And equation (5) means that the unemployment gap (G_t) follows a second-order auto-regressive process, which has a sample mean of zero. The assumption that the unemployment gap follows an autoregressive process is supposed to ensure that - in the absence of shocks - the unemployment rate converges to the structural rate of unemployment. What's more, "specifying the unemployment gap as a process that reverts to a zero mean [...] seems to capture Friedman's (1968) view

[2] As of November 2015 (Autumn 2015 Forecast by the EC), the 'New Keynesian specification' has officially been used by the EC to obtain NAIRU estimates for the following EU countries: Bulgaria, Cyprus, Czech Republic, Denmark, Estonia, Greece, Spain, Finland, France, Hungary, Ireland, Latvia, Poland, Portugal, Romania, Sweden, Slovenia, Slovakia, UK. For the other EU countries, the EC still uses "the so-called traditional Keynesian Phillips (TKP) curve based on static or adaptive expectation assumptions [, which relates] a positive unemployment gap $(u_t - u_t^*)$ with a fall in the change of the growth rate of nominal unit labor cost $(\Delta^2 nulc_t)$ (and vice versa)" (European Commission, 2014, p. 22).

that the unemployment rate cannot be kept away indefinitely from the natural rate [of unemployment]" (Laubach, 2001, p. 221).

The time-path of the NAIRU is extracted from the information contained in the measurement equations by employing the Kalman filter recursions. As the true values of the unobserved components - including the unemployment gap and the NAIRU - are unknown, the Kalman filter provides an algorithm to finding estimates for the unobservables (see Durbin and Koopman, 2012, p. 85ff.).

For the purpose of this paper, it is important to note that neither the variables capturing labor market structures (such as employment protection legislation, unemployment benefits, tax wedge, trade union density etc.) nor the non-structural variables (such as capital accumulation or the long-term interest rate), which might have an impact on the labor market, are included in the model. Nevertheless, the assumption that the NAIRU does eventually represent structural aspects and rigidities in labor markets manifests itself in the EC's treatment of the subject (Orlandi, 2012; European Commission, 2013, 2014; Lendvai *et al.*, 2015).

Whether the NAIRU is determined by structural factors is most crucial when it comes to estimating potential output, which is basically derived from a production function approach making use of empirical data in conjunction with Kalman-filtered estimates for NAIRU (as explained above) and total factor productivity (TFP), where the rationale for filtering the latter is basically to smooth out cyclical variances in productivity growth, given a measure of factor utilization. The conceptual idea behind 'potential output' is to denote a hypothetical level of output at which all production factors would be employed at non-inflationary levels (Havik *et al.*, 2014). In this context, the output gap is used as an indicator for the position of an economy in the business cycle: A positive output gap is said to indicate an over-heating economy, a negative output gap signals underutilization of economic resources. Hence, if there is no discrepancy between actual output and potential output, the *output gap* is zero.

The EC's production function approach is based on the following Cobb-Douglas production function:

$$YPOT_t = L_t^{\alpha} * K_t^{1-\alpha} * TFP_t \tag{6}$$

where $YPOT_t$ is potential output, L_t is the contribution of labor supply to potential output, K_t is the contribution of the capital stock to potential output, and TFP_t is total factor productivity. α and $(1-\alpha)$ are the constant output elasticities of labor and capital, respectively (Havik *et al.*, 2014, p.10).[3]

Since our focus is on the NAIRU, we look more specifically at the estimation of the

[3] The EC assumes that the output-elasticities of labor and capital are equal to 0.65 and 0.35, respectively: "The same Cobb-Douglas specification is assumed for all countries, with the mean wage share for the EU15 over the period 1960-2003 being used as guidance for the estimate of the output elasticity of labor, which would give a value of .63 for α for all Member States and, by definition, .37 for the output elasticity of capital [...] Since these values are close to the conventional mean values of 0.65 and 0.35, the latter are imposed for all countries." (Havik *et al.*, 2014, p. 10)

labor component L_t, which crucially depends on NAIRU estimates:

$$L_t = [POPW_t * PARTS_t * (1 - NAIRU_t)] * HOURST_t \tag{7}$$

where $POPW_t$ is population of working age, $PARTS_t$ is the smoothed labor force participation rate, $NAIRU_t$ is the non-accelerating wage inflation rate of unemployment and $HOURST_t$ is the trend of average hours worked (Havik *et al.*, 2014, p. 14). $PARTS_t$ and $HOURST_t$ are detrended variables; they are calculated by using the Hodrick-Prescott-Filter.[4] It can be seen that potential employment is equal to the labor force – obtained as the product of $POPW_t$ and $PARTS_t$ – times (1-$NAIRU_t$). In other words, estimates of the NAIRU are central to constructing estimates of potential output.[5]

We will now use a replication of the EC's model for estimating the NAIRU and potential output to show how changes in the NAIRU have a direct impact on the scope and evaluation of fiscal policy. The structural budget balance, which is defined as the cyclically-adjusted budget balance, corrected for one-time and temporary effects (e.g. costs related to bailing-out financial institutions), is given by:

$$SB_t = FB_t - \epsilon_t OG_t - OE_t \tag{8}$$

where SB_t is the structural budget balance; FB_t is the reported fiscal balance (defined as government revenues minus government expenditures relative to nominal GDP); ϵ_t is an estimate for the budgetary semi-elasticity, measuring the reaction of the fiscal balance to the output gap (OG_t); and OE_t are one-off and temporary effects (Mourre *et al.*, 2014).

Table 1 illustrates the impact of changes in the NAIRU on potential output and the structural budget balance by using Spain as an example. The EC's official Spanish NAIRU estimate in Autumn 2015 for the year 2015 was 18.5%. In the production function methodology, this NAIRU estimate corresponds to potential output of €1114.8 billion, an output gap of -3.9% and a structural budget balance of -2.5%. Holding everything else constant and assuming that the NAIRU in 2015 would have been estimated to be 1 percentage point lower, we find that potential output rises to €1123.7 billion, an increase of about 0.8% relative to the official estimate. As a consequence, the negative output gap is substantially larger than in the baseline scenario (-4.7% compared to -3.9%), which translates into a decrease in the structural deficit from -2.9% to -2.1% (column 2). The differences are even more pronounced when we assume the Spanish NAIRU in 2015 to be 2.5pp. lower than according to the initial estimates (column 3). Similarly, we can illustrate that upward revisions in the NAIRU compared to the official EC estimates lead to a substantial decrease in potential output going along with an increase in the structural deficit (colums 4 and 5). In other words, the larger (smaller)

[4] The HP filter is a univariate approach to removing the cyclical component of a time series from the trend component (Hodrick and Prescott, 1997). Regarding the basic limitations of the HP filter - with particular emphasis on the so called 'end-point bias' -, see, e.g., Kaiser and Maravall (2001).

[5] While the standard Cobb-Douglas framework is well established, there is still criticism concerning the foundations and the usage of aggregate production functions (e.g. Felipe and McCombie, 2014). This debate, however, is not the focus of this paper.

the estimate of the structural component of unemployment, the larger (smaller) the structural component of the fiscal deficit.

The important point is that the structural budget balance is the central control indicator in the EU's fiscal regulation framework. Crucially, medium-term budgetary objectives (MTOs) for EU countries are defined in terms of the structural budget balance (e.g Ecfin, 2013; Tereanu *et al.*, 2014). In cases where member countries deviate from their MTO, they have to conform to the rules of the Stability and Growth Pact, which require an improvement of the structural budget balance by 0.5% of nominal GDP per year. Since the reform in 2011, the Stability and Growth Pact also stipulates that deviations from the adjustment path to the MTO are significant when the ex-post improvement in the structural budget balance has not amounted to at least 0.5% of GDP in one year or cumulatively over two years (European Union, 2011). According to the European Fiscal Compact, which came into effect on January 1st 2013, the yearly structural deficit may not exceed 0.5% of nominal GDP. The Fiscal Compact also includes the commitment of member countries to codify its rules in national law, preferably in the form of a constitutional safeguard (Fiscal Compact, 2012). Because of this institutionalization of structural budget balances, an increase in the structural deficit translates into more fiscal consolidation pressure.

Against this background, it is essential whether the NAIRU is a good proxy for structural unemployment; otherwise, its usefulness as a key measure for estimating potential employment would be called into doubt. The empirical section of this paper will econometrically investigate the determinants of the NAIRU in order to shed light on the question: What does the NAIRU, as estimated by the EC, actually (not) measure?

	(1) DATA AMECO	**(2) NAIRU -1pp**	**(3) NAIRU -2.5pp**	**(4) NAIRU +1pp**	**(5) NAIRU +2.5pp**
UNEMP	***22.3***	22.3	22.3	22.3	22.3
NAIRU	***18.5***	17.5	16.0	19.5	21.0
YREAL	***1071.1***	1071.1	1071.1	1071.1	1071.1
YPOT	***1114.8***	1123.7	1136.0	1105.9	1092.4
OG	***-3.9***	-4.7	-5.7	-3.1	-2.0
SB	***-2.5***	-2.1	-1.6	-2.9	-3.6

Table 1: Estimates for Spain in 2015: Changes in NAIRU estimates have an impact on potential output and structural budget balances

Notes. Official AMECO data (column 1) is from the Autumn 2015 forecast of the EC. Output gaps and structural budget balances are measured in % of potential output. The calculations are based on the European Commission's potential output model for calculating structural budget balances (Havik *et al.* (2014); Mourre *et al.* (2014); Planas and Rossi (2015)). All scenarios were estimated by holding everything but the NAIRU estimate constant.
UNEMP, unemployment rate; NAIRU, non-accelerating (wage) inflation rate of unemployment; YREAL, GDP at constant prices; YPOT, potential output at constant prices; OG, output gap in % of potential output; SB, structural budget balance in % of potential output.

3 The determinants of (structural) unemployment in European countries: Literature review

Due to the historical rise in European unemployment from the late 1970s to the 1990s, the literature on the cross-country determinants of (structural) unemployment grew rapidly in the 1990s and in the first half of the 2000s, as researchers were trying to explain changes in observed unemployment (see Table 2). A number of influental studies emphasized the link between labor market rigidities imposed by protective labor market institutions and rising unemployment across Europe (e.g. OECD, 1994; Siebert, 1997; International Monetary Fund, 2003; Belot and van Ours, 2004; Nickell *et al.*, 2005; Bassanini and Duval, 2006). This view and corresponding calls for 'structural labor market reforms' provided the dominant theoretical interpretation of increasing unemployment in Europe supported by "a wide range of analysts and international organizations - including the EC, the Organization for Economic Cooperation and Development (OECD), and the International Monetary Fund (IMF) -, [which] have argued that the causes of high unemployment can be found in labor market institutions." (International Monetary Fund, 2003, p. 129)

However, several empirical studies have shown more recently that the empirical evidence for the view that institutions are at the heart of the European unemployment problem from the 1970s to the 1990s is modest at best, since the underlying correlation lacks robustness with regard to variations in control variables, estimation techniques as well as selected countries and time periods (e.g. Howell *et al.*, 2007; Baccaro and Rei, 2007; Stockhammer and Klär, 2011; Vergeer and Kleinknecht, 2012; Avdagic and Salardi, 2013).

	Data	Dependent variable	LMI variables	Other controls
Nickell (1997)	20 OECD countries (1983-1994). Panel with two 6-year averages	UNEMP	UBR, BD, UnD, EPL, CBC, TW, ALMP	—
Elmeskov *et al.* (1998)	19 OECD countries (1983-1995). Panel (annual)	UNEMP	UBR, UnD, EPL, CBC, TW, ALMP, MW	—
Blanchard and Wolfers (2000)	20 OECD countries (1960-1996). Panel with 5-year averages	UNEMP	UBR, BD, UnD, COORD, TW, ALMP, minimum wages	LTI, TFPS, TOTS, LDS
International Monetary Fund (2003)	20 OECD countries (1960-1998). Dynamic panel (annual)	UNEMP	UBR, EPL, UnD, COORD, TW	LTI, TFPS, TOTS, CBI
Belot and van Ours (2004)	17 OECD countries (1960–1999). Panel with 5-year-averages	UNEMP	UBR, EPL, UnD, CWB	—
Baker *et al.* (2005)	20 OECD countries (1960-1999). Panel with 5-year averages	UNEMP	UBR, BD, UnD, EPL, COORD, ALMP	—
Nickell *et al.* (2005)	20 OECD countries (1961-1995). Dynamic panel (annual)	UNEMP	UBR, BD, UnD, EPL, COORD, TW	LTI, TFPS, LDS, TOTS, money supply
Bassanini and Duval (2006)	21 OECD countries (1982-2003). Dynamic panel (annual)	UNEMP	UBR, BD, EPL, UnD, COORD, ALMP; PMR	LTI, TFPS, TOTS, LDS
Palacio-Vera *et al.* (2006)	USA 1964:2-2003:1. Time series	NAIRU (OECD)	—	ACCU, TOTS
Arestis *et al.* (2007)	9 OECD countries (quarterly data, max. 1979-2002). Time series	UNEMP	UBR, strike activity	ACCU
Baccaro and Rei (2007)	18 OECD countries (1960-1998). Dynamic panel; Panel with 5-year-averages	UNEMP	UBR, BD, UnD, EPL, COORD, TW	LTI, TFPS, TOTS, LDS
Bertola *et al.* (2007)	20 OECD countries (1960-1996). Panel with 5-year averages	Employment rate	UBR, BD, UnD, EPL, COORD, ALMP	LTI, TFPS, LDS
Gianella *et al.* (2008)	19 OECD countries (1978-2002). Panel (annual)	NAIRU (OECD)	TW, PMR, UBR, UnD	LTI
Stockhammer and Klär (2011)	20 OECD countries (1983–2003; 1960–1999); Panel with 5-year-averages	UNEMP	UBR, BD, UnD, EPL, TW, COORD, CBC, PMR	TOTS, ACCU, TFPS, LTI, LDS
Orlandi (2012)	13 EU countries (1985–2009). Panel (annual)	NAIRU (EC)	UBR, TW, UnD, ALMP	TFP growth rate, LTI, HBOOM
Vergeer and Kleinknecht (2012)	20 OECD countries (1961-1995). Dynamic panel (annual)	UNEMP	UBR, BD, UD, EPL, COORD, TW	LTI, TFPS, LDS, TOTS, money supply
Avdagic and Salardi (2013)	32 EU and OECD countries (1980–2009). Panel (annual)	UNEMP	UBR, EPL, TW, COORD, UnD	TOTS, LTI, CBI
European Commission (2013)	15 EU Countries (1985–2008). Panel (annual)	NAIRU (EC)	TW, PLM, ALMP, SMI, MEI	TFP growth rate, HBOOM
Flaig and Rottmann (2013)	19 OECD countries (1960–2000). Panel (annual)	UNEMP	EPL, UnD, UBR, CWB, TW	—
Stockhammer *et al.* (2014)	12 OECD countries (2007–2011). Panel (annual)	UNEMP	EPL, ALMP, MW, UnD, GRR	LTI, HBOOM, ACCU

Table 2: Literature review: Selected empirical studies on the determinants of (structural) unemployment

Notes: ACCU, capital accumulation; ALMP, active labor market policy; BD, benefit duration; CBC, collective bargaining coverage; CBI, Central Bank Independence index; COORD, wage bargaining coordination; CWB, centralization of wage bargaining; EPL, employment protection legislation; HBOOM, proxy for boom-bust patterns in housing; LMI, labor market institution; LDS, labor demand shock; LTI, long-term real interest rate; MEI, Matching efficiency indicator; MW, minimum wage; PLM, passive labor market policies; PMR, product market regulation; SMI, skill mismatch indicator; TFPS, deviation of total factor productivity from its trend; TOTS, terms of trade shock; TW, tax wedge; UnD, trade union density; UBR, unemploment benefit replacement rate

The focus in the empirical panel data literature is to explain broad movements in unemployment across OECD countries by shifts in labor market institutions (LMIs) such as trade union density, employment protection legislation, unemployment benefit replacement rate, tax wedge, active labor market policies, minimum wages etc. (see Table 2). As some studies had found no 'meaningful relationship between [the] OECD measure of labor market deregulation and shifts in the NAIRU" (Baker *et al.*, 2005, p. 107), researchers began to include additional control variables representing alternative explanations for the evolution of (structural) unemployment. Blanchard and Wolfers (2000), for instance, control for 'macroeconomic shocks' such as changes in the long-term interest rate, deviations from the trend in total factor productivity growth and shifts in labor demand, emphasizing the link between these shocks and labor market institutions.

Stockhammer and Klär (2011) regard investment as the most crucial variable in explaining unemployment; hence, they include measures of capital accumulation in their regressions. Bassanini and Duval (2006), among others, include a terms of trade shock variable in their regressions, since a change in the terms of trade is assumed to affect domestic unemployment: Whenever a country's terms of trade improve (deteriorate), this implies that for every unit of export sold, this country can purchase more (less) units of imported goods; when imports become less (more) attractive, domestic employment is affected positively (negatively). Finally, Orlandi (2012) introduces another essential control variable, as he considers a proxy for boom-bust-patterns in housing markets. This modification aims to empirically scrutinize the assertion that 'non-structural' factors do not affect 'structural' unemployment at all and, indeed, he finds that in some instances such 'non-structural factors' are "the main drivers of NAWRU developments" (Orlandi, 2012, p. 26).

However, a shortcoming of all major empirical studies on the econometric determinants of unemployment in OECD countries making use of panel data (see Table 2) is that they are characterized by at least one of the following two shortcomings: First, neglecting the role of capital accumulation and investment, the impact of boom-bust patterns related to housing and other macroeconomic developments, like changes in the real interest rate and the terms of trade; second, including only few institutional labor market variables or not considering this aspect at all. Moreover, there are only three studies which have already looked at the determinants of Kalman-filtered NAIRU estimates across several OECD countries, while all the other papers use observed (and in some cases smoothed) unemployment rates as their preferred dependent variable. The relevant papers by Orlandi (2012), the European Commission (2013) and Gianella *et al.* (2008), however, are also incomplete in the sense that they fail to account for the possibility of relevant alternative explanations for the evolution of NAIRU estimates. Our paper closes this gap by analyzing the role of standard labor market variables in explaining the evolution of the EC's NAIRU estimates, while also controlling for a comprehensive set of variables capturing alternative hypotheses with regard to the determinants of the NAIRU.

While the debate on the causes and evolution of European unemployment is again in full swing (e.g. Arpaia *et al.*, 2014; European Central Bank, 2015), in what follows this paper provides an empirical contribution to this debate by econometrically assessing the validity of widely used NAIRU measures for 'structural unemployment' in European countries.

4 Basic econometric strategy and data

The empirical part of this paper analyzes the econometric determinants of the EC's NAIRU estimates. For this purpose, we identified a comprehensive set of explanatory variables covering the basic theoretical and empirical rationales employed in past work and composed a corresponding time-series cross-section data set of 14 countries,[6] for which the complete set of the relevant data could be retrieved. We derive two main specifications from this data: First, we analyze a long-term baseline model based on data for the time period 1985-2011, which covers 11 European OECD countries. Second, we provide an alternative baseline specification focusing on a more recent period (2001-2012). Aside from data considerations – the short term sample allows for the inclusion of 14 countries and two additional LMI variables –, this second specification is motivated by the specific temporal settings, which makes it possible to focus on (a) the euro-era and (b) the run-up and aftermath of the financial crisis.

Our data set enables us to go beyond past contributions on the subject in at least three dimensions: First, we study factors explaining the EC's NAIRU estimates, while nearly all other comparable empirical papers analyze the determinants of observed actual unemployment rates. Second, the time frame of our data set is longer than in comparable studies (Gianella *et al.*, 2008; Orlandi, 2012; European Commission, 2013). In particular, we go beyond past work by including data on the period after the financial crisis of 2008/2009. Third, we look at a more diverse set of potential explanatory variables as compared to past studies. Specifically, we combine data on labor market institutions as provided by the OECD with additional explanatory variables in order to account for alternative hypotheses regarding the evolution of the NAIRU.

The baseline model uses the official NAIRU estimates from the EC's Autumn 2015 forecast as the dependent variable ($NAIRU_{\mathrm{t,\ i}}$). The regression equation has the following form:

$$NAIRU_{\mathrm{t,\ i}} = \beta LMI_{t,i} + \gamma C_{t,i} + \delta_1 FE_{\mathrm{i}} + \delta_2 FE_{\mathrm{t}} + \epsilon_{\mathrm{i,t}}$$

where β represents a vector of regression coefficients related to different structural labor market indicators ($LMI_{\mathrm{t,\ i}}$), while γ is a set of regression coefficients covering other explanatory factors for the NAIRU used in past works ($C_{\mathrm{t,\ i}}$), which will be introduced in Table 3 below. We also introduce country-fixed effects ($\delta_1\mathrm{FE}_i$) to account for unmeasurable, time-invariant country-specific characteristics that may influence the NAIRU as

[6]This group of 14 countries includes: Austria, Belgium, Czech Republic, Denmark, Finland, France, Germany, Ireland, Netherlands, Poland, Portugal, Slovak Republic, Spain, Sweden. Six other countries – Estonia, Greece, Italy, Luxembourg, Slovenia, United Kingdom – have been exluded from the analysis due to data limitations, which are most pronounced in the context of institutional labor market variables.

well as period-fixed effects ($\delta_2\text{FE}_t$) to capture time-varying shocks affecting all countries. $\epsilon_{i,t}$ represents the stochastic residual.

Table 3 provides a detailed overview of the variables included in our data set. Our data on structural labor market indicators ($LMI_{\text{t, j}}$) comprises six standard labor market variables obtained from the OECD's data base: employment protection legislation (EPL), expenditures on active labor market policies (ALMP)[7], trade union density (UnD), unemployment benefit replacement rate (UBR and UBR2)[8], tax wedge (TW) and minimum wage (MW). Variables related to alternative explanations of (structural) unemployment are collected in $C_{\text{t, j}}$ and include the following data: First, we introduce an indicator covering changes in the capital stock (following Stockhammer and Klär, 2011). Capital accumulation (ACCU) in this sense is defined as the ratio of real gross fixed capital formation to the real net capital stock. Second, we employ a proxy for boom-bust-patterns related to the housing market (HBOOM); it is defined as the yearly deviation of the ratio of employment in the construction sector to total employment from its mean – as in Orlandi (2012) . Additionally, we include the annual growth rate in total factor productivity (TFP), a variable for terms of trade shocks (TOTS) and the long-term real interest rate (LTI).

According to Nickell (1998) and other authors who emphasize the role of labor market institutions when it comes to explaining the evolution of (structural) unemployment, UnD, UBR, MW and TW are all expected to have a positive sign, i.e. to be positively associated with (structural) unemployment. The general reasoning is that labor market institutions improve the bargaining position of workers and/or reduce the willingness and capacity of unemployed workers to put downward pressure on wages, which causes labor market rigidities leading to an increase in unemployment.

In contrast, ALMP should have a negative sign, as active labor market policies are expected to increase matching efficiency and, hence, dampen labor market rigidity (e.g. Arpaia *et al.*, 2014). The expected empirical effects of EPL, however, are theoretically ambigous. On one hand, EPL will dampen job creation according to the standard model, because employers are reluctant to hire them due to the fear that they cannot be laid off easily; on the other hand, stricter EPL also increases job retention, as employers lay off fewer employees during economic downturns. What's more, stronger EPL could encourage investments in the training of employees as well as innovation on the firm-level (Zhou *et al.*, 2011), thereby potentially increasing productivity. The effects of EPL are, therefore, ex ante ambiguous (Avdagic and Salardi, 2013).

[7] In this case we use the ratio of ALMP expenditures (as provided by the OECD) to the unemployment rate to account for the fact that ALMP expenditures rise and decrease with current unemployment rates.

[8] For the period 2001-2012, we use OECD data on *net* replacement rates (UBR2). However, as those data are only available until 2001, we have to use *gross* replacements rates for the period 1985-2011 (UBR). The OECD's gross replacement rate data is only available for every second year; therefore, it was interpolated for the missing years. Two separate time series of gross replacement rates were chained. The first series ranges from 1961 to 2005 and is based on Average Production Worker wages; the second time series ranges from 2005 to 2011 and is based on Average Worker wages.

	Data description	Data source
NAIRU	Non-accelerating wage inflation rate of unemployment	AMECO (Autumn 2015 issue)
$LMI_{\text{t, j}}$		
EPL	Strictness of employment protection, individual and collective dismissals (regular contracts)	OECD (December 2nd 2015)
ALMP	Public expenditure and participant stocks on LMP (in % of nominal GDP)	OECD (December 2nd 2015)
UnD	Trade union density	OECD (December 2nd 2015)
UBR	Gross unemployment benefit replacement rate	OECD (December 2nd 2015)
UBR2	Net unemployment benefit replacement rate	OECD (December 2nd 2015)
TW	Average tax wedge (Single person at 100% of average earnings, no child)	OECD (December 2nd 2015)
MW	Real minimum wages (In 2014 constant prices at 2014 USD PPPs)	OECD (December 2nd 2015)
$C_{\text{t, j}}$		
ACCU	Real gross fixed capital formation / real net capital stock	AMECO (Autumn 2015 issue)
HBOOM	Deviation of the ratio of employment in the construction sector to total employment in all domestic industries from its mean	AMECO (Autumn 2015 issue)
LTI	Real long-term interest rates	AMECO (Autumn 2015 issue)
TFP	Yearly growth rate in Total Factor Productivity	AMECO (Autumn 2015 issue)
TOTS	Yearly growth rate in terms of trade index	OECD (December 22nd 2015)
Data for reduced form NAIRU model and different NAIRU forecast vintages		
UNEMP	Unemployment rate	AMECO (Autumn 2015 issue)
ΔINFL	Change in the growth rate of the harmonized consumer price index	IMF World Economic Outlook (October 2015)
NAIRU2014	Non-accelerating wage inflation rate of unemployment	AMECO (Autumn 2014 issue)
NAIRU2013	Non-accelerating wage inflation rate of unemployment	AMECO (Autumn 2013 issue)

Table 3: Variables and data sources

Stockhammer and Klär (2011) provide an additional perspective by emphasizing the role of capital accumulation as an explanatory factor: A decrease in investment causes unemployment to increase (and vice versa), so that ACCU is expected to have a negative sign. LTI also affects capital accumulation; it should be positively associated with unemployment, as an increase in real interest rates is likely to lead to lower aggregate demand, which increases unemployment (e.g. Baker *et al.*, 2005). Orlandi (2012) controls for LTI, but not for ACCU; however, he introduces an additional variable (HBOOM) in his analysis to assess the impact of "severe housing boom-bust effects" (Orlandi, 2012, p. 10). Although from a textbook perspective such 'boom-bust effects' are of a cyclical, transitory nature and should not affect the NAIRU, Orlandi nonetheless posits a negative relationship between HBOOM and NAIRU estimates. According to Blanchard and Wolfers (2000), TFP is expected to have a negative sign, as a decline in TFP growth will cause structural unemployment to increase. Finally, TOTS is a measure for terms of trade shocks, where an improvement in the terms of trade implies that imports become relatively cheaper. Hence, the upward-pressure on wages induced by import-prices is reduced (Bassanini and Duval, 2006, e.g.). It follows that a positive (negative) terms of trade shock is expected to lower (increase) unemployment.

In order to identify a suitable estimation approach for running our regressions, we tested for non-stationarity by running panel unit root tests (Choi, 2001) on the country series for NAWRU, the LMI variables and the additional controls ACCU, HBUB, LTI, TFP and TOTS. For the time period 1985-2011, the null hypothesis that all country series contain a unit root can be rejected for all variables but UnD, EPL, ALMP and LTI. Against the background of these results from the panel unit root tests, we also implemented the test for co-integration proposed by Maddala and Wu (1999), where the null hypothesis is the presence of a unit root in the residuals, i.e. no co-integration amongst the variables. The Maddala-Wu test results signal that estimating our proposed model in levels is appropriate, since the test rejects the null hypothesis of no cointegration at the 1% level, implying that standard OLS and Fixed Effects estimators are consistent.

To ensure robustness of the results, our estimation strategy for analyzing the econometric determinants of the EC's NAIRU estimates is based on two different estimation strategies. In what follows, our preferred estimation technique is to use ordinary least squares (OLS) with panel-corrected standard errors (PCSE), where we include both country- and period-fixed effects. According to Beck and Katz (1995), the OLS-PCSE procedure is well-suited for time-series cross-section models such as ours, where the number of years covered is not much larger than the number of countries in the cross-sectional dimension of the data. The main reason for the superior performance of the OLS-PCSE estimation strategy – compared to the Parks estimator and other Feasible Generalized Least Squares estimators regularly used in the relevant empirical literature – is that the method proposed by Beck and Katz (1995) is well-suited to adressing cross-section heteroskedasticity and autocorrelation in the residuals. Since these two properties are often characteristic of time-series cross-sectional data, the OLS-PCSE estimation strategy helps to avoid overconfidence in standard errors, which is often attributed to the empirical literature on the determinants of unemployment in Europe (Vergeer and Kleinknecht, 2012). Finally, it should be added that this estimation strategy is not an entirely new approach; in fact, using a fixed effects panel estimator in levels is a common estimation technique in recent empirical research on the determinants of (structural) unemployment (e.g. Flaig and Rottmann, 2013), with some authors also following the OLS-PCSE estimation and correction procedure as implemented in this paper (Orlandi, 2012; Avdagic and Salardi, 2013).

This preferred estimation approach is complemented by using a first difference estimator applied to annual data and 5-year-data averages, respectively. In accordance with Baccaro and Rei (2007) we find that using first differences of 5-year-average-data removes the positive autocorrelation in the residuals, which is characteristic of our baseline regression results. Aside from this econometric justification, the economic rationale for using 5-year-averages has two aspects: First, it takes into account that labor market institutions only change slowly. Second, it dampens possible effects of business cycle fluctuations on (structural) unemployment, which should allow for more reliable causal interpretations. The obvious drawback from using averages, however, is a loss of information as contained in the data, which makes it especially difficult to trace short-term effects between our explanatory variables and NAIRU estimates, as well as a drastic reduction of observations, which lowers the statistical power of the test.

Against this backdrop, our preferred estimation strategy is to use annual data in levels in a time-series cross-section model with OLS-PCSE, while our alternative estimation strategy based on first-differences of either annual data or 5-year averages is used primarily as an additional tool examining the robustness of single relationships between the explanatory variables and the NAIRU estimates.

5 Econometric baseline results

The econometric baseline results for 11 European OECD countries over the time period 1985-2011 from six different models are shown in Table 4. In the first column, we regress the EC's NAIRU estimates on four instititutional labor market indicators (EPL, ALMP, UnD, UBR); in addition, we control for TFP and TOTS. Arguably, this specification leaves ample scope for the institutional variables to explain the variation in the dependent variable. The regression coefficients represent the impact of a 1 unit increase in the respective explanatory variable on the NAIRU (in percentage points). For example, an increase in the unemployment benefit replacement rate (UBR) by 10 percentage points increases the NAIRU by about 0.9 percentage points. Standard errors of the fixed effects models shown in Table 4 are PCSE-corrected standard errors. As both Durbin-Watson (DW) and Breusch-Godfrey (BG) tests on autocorrelation indicate positive serial correlation in the residuals, the PCSE procedure is a sensible tool to account for this data characteristic in our fixed effects models.

In model 1, all coefficients of the institutional variables are signed as expected in the standard literature on the determinants of structural unemployment. However, only ALMP is statistically significant at the 5% level, while UBR is weakly significant using a 90% confidence interval. The adjusted R^2 suggests that the regressors are merely able to explain about 20% of the variation in the EC's NAIRU estimates. In brief, the results from column 1 suggest that we ought to reject the hypothesis that NAIRU estimates can be exclusively explained by differences in labor market institutions and productivity growth.

In model (2), we therefore introduce capital accumulation and the long-term real interest rate to account for alternative hypotheses aiming to explain the evolution of the EC's NAIRU estimates. The introduction of those two additional variables leads to a tripling of the adjusted R^2, which changes to 58%. LTI is positively signed (but insignificant), while ACCU - as expected in the relevant literature (Stockhammer and Klär, 2011) - is negatively signed and strongly significant, with the coefficient implying that an increase in the ratio of real gross fixed capital formation to the real net capital stock by 1 percentage point lowers the NAIRU by 1.5 percentage points. The size of the coefficients of the institutional variables in column (2) changes to varying degrees, while the estimated direction of the effects remains the same. EPL turns weakly significant, while UBR is now significant at the 1% level. In model (3), we again exclude ACCU, but instead introduce our proxy for boom-bust patterns in housing (HBOOM), which is signed as expected and highly significant, suggesting that boom (bust) patterns in housing are associated with decreases (increases) in the NAIRU. It is also notable that the coefficient of LTI in this setup is markedly larger than in column 2 and significant at the 5% level.

	Dependent variable: **NAIRU**					
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS-PCSE	OLS-PCSE	OLS-PCSE	OLS-PCSE	FD	FD
ACCU		−1.509***		−1.327***	−0.226***	−0.721***
		(0.177)		(0.233)	(0.071)	(0.261)
HBOOM			−0.998***	−0.242	−0.289***	−0.565*
			(0.187)	(0.196)	(0.075)	(0.288)
LTI		0.071	0.238**	0.064	0.032**	0.064
		(0.060)	(0.094)	(0.063)	(0.016)	(0.112)
EPL	0.485	1.660*	−0.134	1.391	0.088	1.681**
	(1.782)	(0.936)	(1.204)	(0.904)	(0.274)	(0.726)
ALMP	−0.050**	−0.029**	−0.037**	−0.029**	−0.004	−0.027**
	(0.025)	(0.013)	(0.017)	(0.013)	(0.004)	(0.012)
UnD	0.100	0.056	0.092	0.058	0.055***	0.100
	(0.091)	(0.048)	(0.065)	(0.047)	(0.020)	(0.060)
UBR	0.089*	0.072***	0.096**	0.080***	0.016*	0.102***
	(0.053)	(0.025)	(0.039)	(0.025)	(0.009)	(0.036)
TFP	0.015	−0.104	−0.229***	−0.145**	0.001	−0.417*
	(0.088)	(0.067)	(0.085)	(0.069)	(0.010)	(0.241)
TOTS	−0.079	0.008	−0.002	−0.006	0.004	−0.078
	(0.084)	(0.062)	(0.071)	(0.060)	(0.009)	(0.180)
Constant					0.064***	0.116
					(0.022)	(0.250)
Countries	11	11	11	11	11	11
Time periods	27	27	27	27	26	4
Observations	297	297	297	297	286	44
Adjusted R^2	0.195	0.582	0.463	0.586	0.323	0.553
Country FE	√	√	√	√		
Period FE	√	√	√	√		
DW test	0.182	0.448	0.382	0.476	0.382	1.900

*p<0.1; **p<0.05; ***p<0.01

Table 4: Results for 1985-2011

Notes.
(1)-(4) OLS-PCSE. Standard errors in brackets () corrected for autocorrelation in residuals. Cross-section and Year Fixed Effects. Standard errors are illustrated in brackets ().
(5) First difference estimator. Heteroskedasticity-robust standard errors.
(6) First difference estimator, five-year-averages. Heteroskedasticity-robust standard errors.
Country group in all specifications: Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Netherlands, Portugal, Spain, Sweden.
DW test denotes the Durbin-Watson test statistic on autocorrelation in the residuals.
NAIRU, non-accelerating (wage) inflation rate; ACCU, capital accumulation; HBOOM, housing boom/bust proxy; LTI, long-term real interest rate; EPL, employment protection legislation; ALMP, active labor market policies; UnD, trade union density; UBR, gross unemployment benefit replacement rate; TFP, total factor productivity; TOTS, terms of trade shock.

However, as soon as we include all our regressors at once in column 4, LTI and HBOOM turn insignificant, while the coefficient of ACCU remains negative, large and highly significant, which supports the earlier finding from model 3 that capital accumulation plays an important part in explaining NAIRU estimates in our data set of European OECD countries. According to model 4, an increase in the ratio of capital formation to the capital stock by 1 percentage point lowers the NAIRU by approximately 1.3 percentage points, while a 10 percentage point increase in UBR increases the NAIRU by 0.8 percentage points.

One possible issue with model 4 could be that the inclusion of fixed effects has an impact on the size and significance of the LMI coefficients. In order to investigate this issue, we also ran regressions with country-fixed effects only. We find that the LMI coefficients and their significance do not change markedly when we exclude period-fixed effects, while the coefficient of HBOOM nearly doubles to -0.45 and turns significant at the 5% level; ACCU retains its significance at the 1% level.

In model 5, we employ a First Difference estimator to the annual data (with robust standard errors). Notably, all institutional variables are again signed as expected, but remain statistically insignificant with the exception of UBR (weakly significant) and UnD (strongly significant). In this specification, capital accumulation, the housing boom/bust proxy and the long-term interest rate have a significant effect on NAIRU estimates. Finally, model 6 follows the strategy preferred by Baccaro and Rei (2007), i.e. deploying the First Difference estimator after calculating 5-year-averages for all time series. Regarding the institutional variables, model 6 finds EPL, ALMP and UBR to be signed as expected as well as statistically significant (at different levels of confidence). However, the major finding that capital accumulation and housing booms and busts are controls that ought not to be omitted when trying to explain the EC's NAIRU estimates, is also retained in this final specification.

	Dependent variable: **NAIRU**					
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS-PCSE	OLS-PCSE	OLS-PCSE	OLS-PCSE	FD	OLS-PCSE
ACCU		-0.858^{***}		-0.627^{***}	-0.139^{**}	-0.479^{**}
		(0.161)		(0.186)	(0.070)	(0.224)
HBOOM			-0.559^{***}	-0.303^{*}	-0.353^{***}	-0.262
			(0.139)	(0.161)	(0.080)	(0.184)
LTI		0.211^{***}	0.239^{***}	0.188^{***}	0.031	0.244^{***}
		(0.057)	(0.062)	(0.058)	(0.019)	(0.071)
EPL	-2.480^{*}	0.591	0.311	0.529	0.046	0.147
	(1.401)	(0.973)	(0.950)	(0.914)	(0.314)	(1.063)
ALMP	-0.142^{***}	-0.064^{***}	-0.084^{***}	-0.064^{***}	-0.006	-0.079^{***}
	(0.029)	(0.019)	(0.020)	(0.018)	(0.008)	(0.029)
UnD	0.134	0.064	0.033	0.043	-0.005	0.115^{*}
	(0.096)	(0.056)	(0.068)	(0.057)	(0.021)	(0.066)
UBR2	0.182^{***}	0.111^{***}	0.124^{***}	0.113^{***}	0.016^{**}	0.146^{***}
	(0.038)	(0.021)	(0.026)	(0.022)	(0.008)	(0.034)
TW	0.315^{**}	0.032	0.048	-0.017	-0.001	0.042
	(0.132)	(0.095)	(0.104)	(0.096)	(0.028)	(0.123)
MW						-0.00003
						(0.0002)
TFP	0.110	0.051	-0.016	0.009	0.008	0.036
	(0.068)	(0.057)	(0.057)	(0.055)	(0.009)	(0.085)
TOTS	0.026	0.183^{***}	0.180^{***}	0.173^{***}	0.021	0.124
	(0.076)	(0.058)	(0.064)	(0.057)	(0.013)	(0.079)
Constant					0.020	
					(0.028)	
Countries	14	14	14	14	14	9
Time periods	12	12	12	12	12	12
Observations	168	168	168	168	154	108
Adjusted R^2	0.479	0.627	0.612	0.633	0.339	0.595
Country FE	√	√	√	√		√
Period FE	√	√	√	√		√
DW test	0.851	0.884	0.769	0.795	0.684	0.995

*p<0.1; **p<0.05; ***p<0.01

Table 5: Results for 2001-2012

Notes.
(1)-(4), (6) OLS-PCSE. Standard errors in brackets () corrected for autocorrelation in residuals.. Cross-section and Year Fixed Effects.
(5) First difference estimator. Heteroskedasticity-robust standard errors.
Country group for specifications (1)-(5): Austria, Belgium, Czech Republic, Denmark, Finland, France, Germany, Ireland, Netherlands, Poland, Portugal, Slovak Republic, Spain, Sweden
Due to missing MW data, specification (6) excludes Austria, Denmark, Finland, Germany and Sweden.
DW test denotes the Durbin-Watson test statistic on autocorrelation in the residuals.
NAIRU, non-accelerating (wage) inflation rate; ACCU, capital accumulation; HBOOM, housing boom/bust proxy; LTI, long-term real interest rate; EPL, employment protection legislation; ALMP, active labor market policies; UnD, trade union density; UBR2, net unemployment benefit replacement rate; TW, tax wedge; MW, minimum wage; TFP, total factor productivity; TOTS, terms of trade shock.

Table 5 illustrates the baseline results for the time period 2001-2012, where all model specifications with the exception of model 6 are the same as in Table 3. Looking at the institutional variables, we again find that – with very few exceptions – all LMIs are signed as expected across the different model specifications. As in the time period 1985-2011, ALMP and UBR are again the only significant LMI variables. We also support the major finding from the longer time period that ACCU plays an important part in explaining the NAIRU: In all columns, ACCU is at least significant at the 5% level. LTI has a larger coefficient and seems to play a somewhat stronger role than over 1985-2011, as it is highly significant in nearly all of the relevant models. Moreover, HBOOM is also again signed as expected and statistically significant in the majority of scenarios. Summing up, running regressions on the shorter time period of 2001-2012 – for which data availability for LMIs has improved – supports our baseline findings from 1985-2011. This suggests that the EC's implicit assumption that NAIRU estimates gained by detrending the unemployment rate are a good proxy for 'structural unemployment' does not hold. On the contrary, most institutional variables are either statistically insignificant or their significance is sensitive to the model specification, while cyclical factors - especially capital accumulation - play a prominent role in explaining NAIRU estimates.

6 Robustness checks

To assess the sensitivity of the baseline results, this section discusses several robustness checks: Specifically, we analyze the impact of variations in the country group, introduce lag specifications, consider interaction terms and, finally, implement variations in the dependent variable.

The first sensitivity test consists of checking whether our overall baseline results are driven by outlier countries. Therefore, we varied the country group by excluding one country at a time. The results from this variation allow us to conclude that for both the long period (1985-2011) and the shorter period (2001-2012) neither the size of the coefficients of the explanatory variables nor their statistical significance are markedly affected by including or excluding single countries.[9]

In a second step, we investigated how the introduction of lags affects our regression results. In doing so we use specification (4) from the baseline models as a reference point, as it includes all major variables that proved to be empirically relevant in our past explorations. Table 6 depicts lag specification results for both time periods, where columns (1)-(3) refer to 1985-2011 and columns (4)-(6) depict the results for 2001-2012. In columns (1) and (4) we introduce lags for all the LMI variables to allow for the argument that institutional changes tend to affect the NAIRU with a lag, which could also have an impact on the performance of our alternative explanatory variables. However, this hypothesis is not supported by the regression results, as coefficients and standard errors of the variables ACCU, HBOOM and LTI remain largely unaffected after we introduce LMI lags, while the institutional variables either have a sign that is not in line with

[9]The detailed regression results from varying the country group are available in the online data appendix.

their standard theoretical prediction or they are statistically insignificant. We proceeded by including lags for capital accumulation, the housing boom/bust proxy and the real interest rate in columns (2) and (5) to find out whether these alternative factors impact on the NAIRU with a lag. We confirm the central role of ACCU in explaining the EC's NAIRU estimates, although the ACCU coefficient in column 2 is less negative due to the introduction of the statistically significant ACCU lag. In columns (3) and (6) we include all possible lag terms: both for the LMI variables, and ACCU/HBOOM/LTI; in addition, we also consider lags for TOTS and TFP. The main results from the reference model in the baseline tables, however, still hold: While they underscore the importance of alternative factors - especially ACCU - in driving the NAIRU, the econometric evidence for the role of LMI variables is at best mixed.

A third sensitivity topic are interaction terms, as the econometric literature contains several papers which emphasize that LMIs should be expected to have an effect on (structural) unemployment through their interactions (e.g. International Monetary Fund, 2003; Bassanini and Duval, 2006). In a seminal paper, Blanchard and Wolfers (2000) stress the role of interactions between LMI variables and macroeconomic shocks. A major problem in this literature, however, is that "[t]he theoretical foundation for these interactions is [...] unspecific. For example, the IMF (2003) argues that the effects of different LMI are reinforcing, without specifying ex ante which LMI should interact. This poses a problem for an attempt to statistically evaluate the effects of interactions: since there are numerous potential interactions, the inclined researcher is bound to find some that prove statistically significant." (Stockhammer and Klär, 2011, p. 449).

Nevertheless, we accounted for possible interactions by looking at various interaction specifications. No matter whether we include interactions between LMIs only, interactions among LMIs and the other macroeconomic controls only, or all interactions at once, the result is always that there is no systematic evidence that the effects of different LMI variables are reinforced by their interactions. This leads us to the interpretation that the data do not support the argument that LMI interaction terms are crucial for explaining the EC's NAIRU estimates.[10]

[10] Again, detailed results from introducing interaction terms can be found in the online data appendix.

	Dependent variable: ***NAIRU***					
	(1)	(2)	(3)	(4)	(5)	(6)
Time period	1985-2011	1985-2011	1985-2011	2001-2012	2001-2012	2001-2012
Estimator	OLS-PCSE	OLS-PCSE	OLS-PCSE	OLS-PCSE	OLS-PCSE	OLS-PCSE
EPL	1.093	1.725**	1.337*	0.714	0.332	0.748
	(0.725)	(0.839)	(0.792)	(0.804)	(0.870)	(0.821)
EPL_{t-1}	0.903		0.325	−0.555		−0.704
	(0.662)		(0.685)	(1.148)		(1.107)
ALMP	0.007	−0.027**	−0.009	−0.007	−0.071***	−0.010
	(0.016)	(0.011)	(0.015)	(0.021)	(0.022)	(0.022)
$ALMP_{t-1}$	−0.035**		−0.019	−0.079***		−0.072***
	(0.015)		(0.015)	(0.024)		(0.024)
UnD	−0.237***	0.056	−0.185**	0.016	0.064	0.042
	(0.086)	(0.043)	(0.093)	(0.061)	(0.063)	(0.061)
UnD_{t-1}	0.291***		0.203**	0.052		0.054
	(0.089)		(0.093)	(0.067)		(0.069)
UBR	0.026	0.078***	0.032			
	(0.044)	(0.025)	(0.045)			
UBR_{t-1}	0.062		0.063			
	(0.051)		(0.047)			
UBR2				0.097***	0.118***	0.096***
				(0.023)	(0.024)	(0.024)
$UBR2_{t-1}$				0.051**		0.049**
				(0.020)		(0.021)
TW				−0.151*	−0.071	−0.168**
				(0.078)	(0.100)	(0.076)
TW_{t-1}				0.157*		0.117
				(0.087)		(0.095)
ACCU	−1.468***	−0.778***	−0.668***	−0.637***	−0.492**	−0.578***
	(0.236)	(0.247)	(0.255)	(0.174)	(0.192)	(0.165)
$ACCU_{t-1}$		−0.686**	−0.524**		−0.284	−0.159
		(0.285)	(0.265)		(0.195)	(0.182)
HBOOM	−0.159	0.211	−0.159	−0.209	0.137	0.029
	(0.202)	(0.316)	(0.322)	(0.158)	(0.239)	(0.217)
$HBOOM_{t-1}$		−0.363	−0.166		−0.359*	−0.165
		(0.262)	(0.303)		(0.181)	(0.172)
LTI	0.081	0.107*	0.069	0.163**	0.173***	0.128**
	(0.062)	(0.059)	(0.049)	(0.063)	(0.060)	(0.060)
LTI_{t-1}		0.050	−0.032		0.069	0.103*
		(0.060)	(0.048)		(0.054)	(0.055)
	(0.040)		(0.033)	(0.053)		(0.057)
TFP	−0.140**	−0.221***	−0.188***	0.064	0.018	0.062
	(0.067)	(0.070)	(0.069)	(0.049)	(0.052)	(0.049)
TFP_{t-1}	−0.049		−0.014	0.010		0.004
	(0.062)		(0.050)	(0.049)		(0.051)
TOTS	−0.036	−0.006	−0.132***	0.108*	0.127**	0.067
	(0.048)	(0.059)	(0.037)	(0.060)	(0.056)	(0.059)
$TOTS_{t-1}$	0.037		−0.090***	0.068		0.096*
Countries	11	11	11	14	14	14
Time periods	26	26	26	11	11	11
Observations	286	286	286	154	154	154
Adjusted R^2	0.605	0.602	0.625	0.610	0.617	0.601
Country FE	√	√	√	√	√	√
Period FE	√	√	√	√	√	√
DW test	0.551	0.479	0.515	1.014	0.900	0.963

*p<0.1; **p<0.05; ***p<0.01

Table 6: Lag specifications; results for 1985-2011 and 2001-2012

Notes.
(1)-(6) OLS-PCSE. Standard errors in brackets () corrected for autocorrelation in residuals. Cross-section and Year Fixed Effects.
Country group in specifications (1)-(3): Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Netherlands, Portugal, Spain, Sweden.
In specifications (4)-(6), we additionally include Czech Republic, Poland and Slovak Republic.
DW test denotes the Durbin-Watson test statistic on autocorrelation in the residuals.
NAIRU, non-accelerating (wage) inflation rate; ACCU, capital accumulation; HBOOM, housing boom/bust proxy; LTI, long-term real interest rate; EPL, employment protection legislation; ALMP, active labor market policies; UnD, trade union density; UBR, gross unemployment benefit replacement rate; UBR2, net unemployment benefit replacement rate; TW, tax wedge; MW, minimum wage; TFP, total factor productivity; TOTS, terms of trade shock.
$_{t-1}$ denotes the first lag of the respective variable; e.g., EPL_{t-1} is the first lag of employment protection legislation.

As a fourth and final robustness check, we implemented variations in the dependent variable. As researchers have noted sizeable revisions in the EC's NAIRU estimates since the outbreak of the financial crisis (Cohen-Setton and Valla, 2010; Klaer, 2013), we also obtained NAIRU data from earlier forecast vintages to assess the robustness of our results with respect to a change in measuring the dependent variable. In columns (1) and (2) of table 7, we employ the EC's NAIRU estimates from Autumn 2014 and Autumn 2013 for the time period 1985-2011, respectively. We then proceed with another sensitivity check: In columns (3)-(6) we use the actual unemployment rate as the dependent variable. The change in the inflation rate (ΔINFL) was introduced as an additional control variable to capture a possible trade-off in the Phillips curve relationship between unemployment and inflation - a feature of the reduced form NAIRU models used in the empirical literature on the determinants of unemployment (Nickell, 1997; Stockhammer and Klär, 2011, e.g.). We report the reduced form NAIRU model results for the time period 1985-2011 (column 3) and 2001-2012 (column 5) with country- and period-fixed effects, estimated by OLS-PCSE. Results from the First Difference estimator are shown in columns (4) and (6). In all these variations, it is evident that ACCU and HBOOM are signed as expected, and they are highly significant in all the reduced form NAIRU models. In contrast, ALMP and (partially) UBR are the only LMI variables that are consistently signed as expected and significant across all models.

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable	NAIRU2014	NAIRU2013	UNEMP	UNEMP	UNEMP	UNEMP
Time period	1985-2011	1985-2011	1985-2011	1985-2011	2001-2012	2001-2012
Estimator	OLS-PCSE	OLS-PCSE	OLS-PCSE	FD	OLS-PCSE	FD
ΔINFL			−0.036	−0.042	−0.172**	−0.095***
			(0.051)	(0.026)	(0.077)	(0.033)
ACCU	−1.285***	−1.384***	−1.634***	−1.188***	−1.619***	−1.137***
	(0.222)	(0.240)	(0.237)	(0.133)	(0.361)	(0.128)
HBOOM	−0.294	−0.459**	−0.587**	−0.790***	−0.861***	−0.985***
	(0.193)	(0.201)	(0.240)	(0.165)	(0.291)	(0.172)
LTI	0.073	0.035	0.149*	0.002	0.042	−0.077
	(0.061)	(0.072)	(0.083)	(0.032)	(0.121)	(0.051)
EPL	1.115	1.411*	1.727**	−0.355	0.776	0.614
	(0.880)	(0.844)	(0.749)	(0.702)	(1.560)	(0.621)
ALMP	−0.031***	−0.026**	−0.043***	−0.051***	−0.107***	−0.071***
	(0.012)	(0.012)	(0.012)	(0.009)	(0.040)	(0.018)
UnD	0.018	0.040	0.032	0.140**	0.057	0.108
	(0.044)	(0.044)	(0.042)	(0.070)	(0.118)	(0.082)
UBR	0.068***	0.075***	0.056**	−0.009		
	(0.023)	(0.028)	(0.028)	(0.016)		
TW					−0.001	−0.037
					(0.199)	(0.066)
UBR2					0.104**	−0.001
					(0.044)	(0.012)
TFP	−0.168**	−0.120*	−0.095	0.053***	−0.027	0.103***
	(0.068)	(0.072)	(0.075)	(0.019)	(0.092)	(0.030)
TOTS	−0.027	−0.013	0.042	−0.016	0.072	0.017
	(0.058)	(0.060)	(0.069)	(0.027)	(0.096)	(0.036)
Constant				0.064		0.005
				(0.054)		(0.066)
Countries	11	11	11	11	14	14
Time periods	27	27	27	26	12	11
Observations	297	297	297	286	168	154
Adjusted R^2	0.587	0.610	0.659	0.669	0.652	0.676
Country FE	√	√	√		√	
Period FE	√	√	√		√	
DW test	0.485	0.484	0.541	1.295	0.860	1.384

*p<0.1; **p<0.05; ***p<0.01

Table 7: Results for 1985-2011: Further robustness checks

Notes.
(1)-(3) and (5): OLS-PCSE. Standard errors in brackets () corrected for autocorrelation in residuals. Cross-section and Year Fixed Effects.
(4) and (6): First Difference Estimator (FD).
Country group in specifications (1)-(4): Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Netherlands, Portugal, Spain, Sweden.
In specifications (5)-(6), we additionally include Czech Republic, Poland and Slovak Republic.
DW test denotes the Durbin-Watson test statistic on autocorrelation in the residuals.
NAIRU, non-accelerating (wage) inflation rate (Autumn 2015); NAIRU2014, non-accelerating (wage) inflation rate (Autumn 2014); NAIRU2013, non-accelerating (wage) inflation rate UNEMP (Autumn 2013); unemployment rate; ΔINFL, change in the inflation rate; ACCU, capital accumulation; HBOOM, housing boom/bust proxy; LTI, long-term real interest rate; EPL, employment protection legislation; ALMP, active labor market policies; UnD, trade union density; UBR, gross unemployment benefit replacement rate; UBR2, net unemployment benefit replacement rate; TW, tax wedge; MW, minimum wage; TFP, total factor productivity; TOTS, terms of trade shock.

7 Discussion: The NAIRU in theory, empirics and policy

Our setup for analyzing the econometric determinants of the EC's NAIRU estimates leads to a confrontation between theory and empirics: While the NAIRU is a theoretically postulated concept, which explains structural unemployment by institutional rigidities, its estimation in the particular context is largely devoid of theoretical rationales, but rather follows a Kalman-Filter approach for detrending time-series data. It is, hence, more of an econometric than an economic exercise.

Against this backdrop, our results raise some skepticism with regard to the adequacy of the EC's NAIRU estimates. However, we cannot provide a conclusive answer about whether the poor fit between NAIRU estimates and their supposed structural explanatory variables is due to principal theoretical deficiencies or rather has to be attributed to a sub-optimal performance of the underlying Kalman-filtering techniques for estimating the NAIRU. Nonetheless, our analysis allows for a closer examination of 'what's wrong' with the EC's NAIRU estimates.

According to the econometric findings discussed in the previous sections, we find that the performance of labor market institutions with regard to explaining the EC's NAIRU estimates is moderate at best. In the specifications that we tested, variables such as the tax wedge, union density, employment protection legislation and minimum wages either do not have the sign expected by standard theory or they are statistically insignificant. This finding points to a contradiction with the theoretical framework used by the EC, which implicitly assumes that the NAIRU is a good proxy for structural unemployment, driven by institutional factors. Orlandi (2012) found for 13 EU countries covering the period 1985-2009 that structural labor market indicators provide a good fit for "[the] structural unemployment rate, as measured by the Commission services (i.e. the so-called NAWRU)" (Orlandi, 2012, p. 1). Similarly, Gianella *et al.* (2008) had reported that "the set of structural variables provides a reasonable explanation of [the OECD's Kalman-filtered] NAIRU dynamics over the period 1978-2003" (Gianella *et al.*, 2008, p. 1). In our empirical analysis, we went beyond these earlier studies in many respects. Most crucially, we included additional alternative explanatory factors for the NAIRU and took the years after the financial crisis into account. Our findings are in stark contrast to the assessments by Orlandi (2012) and Gianella *et al.* (2008): Given that institutional variables underperform in our regressions, we conclude that the NAIRU is *not* a good proxy for 'structural unemployment'. This point is reemphasized by the central role that cyclical factors – such as capital accumulation and the housing boom/bust proxy – play in our regressions when it comes to explaining the EC's NAIRU estimates.

Finally, our results provide food for thought regarding more general drawbacks imposed by a 'one-size-fits-all' analytical approach to understanding unemployment in Europe, especially as such a framework, quite naturally, translates into a 'one-size-fits-all' policy approach. With regard to the analytical aspects we should ask which cyclical variables affect the NAIRU estimates of different countries, and, hence, whether NAIRU estimates might also require context-sensitive interpretations, depending on the country under study.[11] In this context, it is remarkable that, although the economic situation of the Eurozone countries exhibits considerable variation, the policy approach suggested by the EC is, nonetheless, quite uniform: 'Structural reforms' which aim at deregulating labor markets are thereby widely recommended, as member countries are urged to lower structural unemployment by supply side reform (Canton *et al.*, 2014).

We argue that a more nuanced analytical approach, which more clearly departs from 'one size fits all', is in order. Such a new approach would have to allow for the incorporation of a more diverse set of facts, e.g. that Germany's competitiveness is rather based on sectoral specialization and strong 'non-price' competitiveness than on flexible labor markets (Carlin *et al.*, 2001; Storm and Naastepad, 2015). Another example would be to consider whether Spain's and Ireland's NAIRU before and after the financial crisis might actually have been pro-cyclically driven by the development of their respective housing markets and the repercussions of the boom-bust cycle in the labor markets, as indicated by the strong relationship between the housing boom/bust proxy and the NAIRU plotted in Figure 1. By considering different structural and cyclical factors that impact on NAIRU estimates in specific countries, a nuanced approach would allow for devising more flexible, adaptive and versatile policy strategies by more effectively taking into account the economic idiosyncracies of individual countries.

Since the NAIRU is used as a proxy for 'structural unemployment' in calculating potential output and structural budget balances in EU member countries, so that it has a direct impact on the scope and evaluation of fiscal policy (see section 2), a framework considering the role of institutional and cyclical factors in driving NAIRU estimates would be superior to the predominant approach preferred by the EC, which implicitly assumes that Kalman-filter estimates of the NAIRU reflect 'purely' structural factors, stripped off any cyclical influences. Our analysis shows that both economists and policymakers have to be cautious in interpreting NAIRU estimates as a useful measure for 'structural unemployment' that can unambiguously be used to assess the contribution of the production factor labor to potential output. On the contrary, our econometric findings suggest that the predominant framework for coordinating fiscal policies in the euro area may be dysfunctional, because it crucially rests on an econometric estimate of the NAIRU that does neither correspond to its key theoretical postulate nor to its political application. Eventually, this poses the risk of using a deficient measure - the output gap - for judging what's 'structural' about fiscal deficits, thereby misinforming policy-making at large.

[11] This observation is also well in line with the variation of fixed effects as estimated by our models. For example, we find for model 4 in Table 4 that the country fixed effects coefficients vary from a minimum of 4.0 in the case of Sweden to a maximum of 18.5 for Spain.

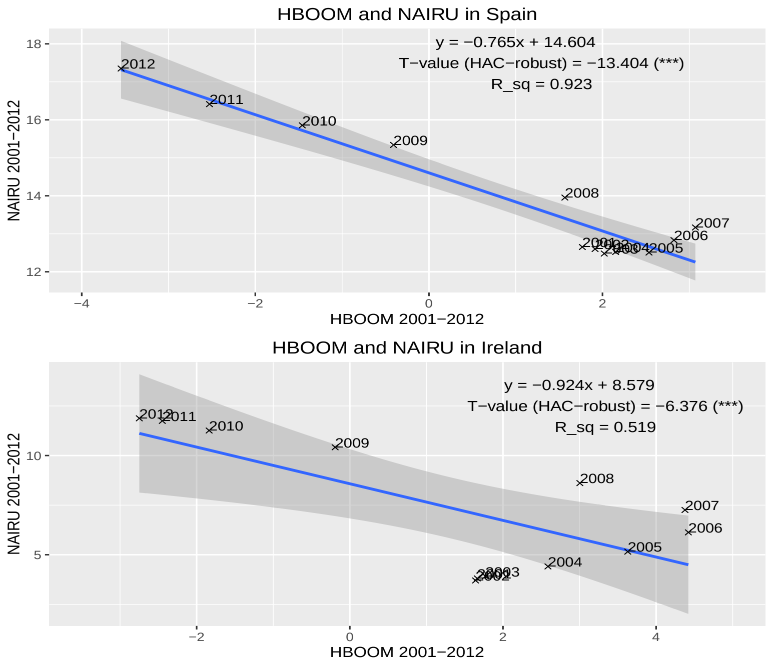

Figure 1: Correlation of HBOOM and NAIRU in Spain and Ireland (2001-2012), respectively.

Data: AMECO (Autumn 2015.); authors' calculations

8 Conclusions

This paper has analyzed the determinants of the European Commission's NAIRU estimates for 14 European OECD countries over the time period 1985-2012. Our main finding is that the NAIRU, as estimated by the EC, is *not* a good proxy for 'structural unemployment': Most indicators of labor market institutions – employment protection legislation, union density, tax wedge and minimum wage – do not explain much; either is their sign inconsistent with the expectation from standard theory, they are statistically insignificant, or their significance is sensitive to the model specification. Only active labor market policies and unemployment benefit replacement rates are consistently signed as expected and significant. The point that NAIRU estimates are not simply driven by institutions is underscored by the finding that cyclical factors – especially capital accumulation and boom-bust patterns in housing markets – are important determinants. This shows that the empirics of the NAIRU are in conflict with the EC's theoretical framework, in which the NAIRU is modeled as the trend component of the unemployment rate, stripped off all cyclical factors.

Our econometric findings are highly relevant for policy making in the EU. First, they point to the fact that increases in the NAIRU cannot simply be attributed to more institutional rigidities with corresponding calls for labor market deregulation to lower 'structural unemployment'. At the same time, they indicate that the causes for a decline in the NAIRU in a specific country are not always to be found in successful labor market reforms: Downward revisions in the NAIRU might also be driven by cyclical factors. Second, our findings show that there is a considerable risk that NAIRU estimates – which are at least partly driven by cyclical factors – misinform fiscal policy making in the EU. The reason is that the NAIRU is used as a proxy for 'structural unemployment' in calculating output gaps as a measure for the position of an economy in the business cycle - an indicator that is then transformed into a judgement on how much of the fiscal deficit is due to structural and cyclical factors, respectively. Accordingly, flawed estimates of the NAIRU as the 'structural unemployment rate' can lead to miscalculations of the size of the structural deficit and inappropriate fiscal policies.

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