Socio-Economic Review (2013) **11**, 739–769 doi:10.1093/ser/mwt009
Advance Access publication May 6, 2013

Tenuous link: labour market institutions and unemployment in advanced and new market economies

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International organizations and mainstream economists have consistently promoted the view that labour market rigidities are responsible for high unemployment, and that wide-ranging institutional deregulation is an appropriate policy response. Yet, as demonstrated by recent literature, the empirical support for the deregulatory view is ambiguous. This paper re-assesses this debate by bringing in new evidence from a larger group of countries, which includes advanced and new market economies. Using new data and paying special attention to the robustness of estimation results, we find rather thin support for the deregulatory view. The sensitivity analysis demonstrates that in most cases the adverse effects of institutions disappear with small changes in the sample or the use of alternative estimators and specifications. The impact of institutions is particularly weak in new market economies, where unemployment is related primarily to macroeconomic factors. Overall, our findings challenge the policy orthodoxy that comprehensive deregulation is the universal solution to unemployment.

Keywords: unemployment, labor market institutions, OECD countries, Central and Eastern Europe
JEL classification: J4 and J48 General Labour Markets and Public Policy, P16 Political Economy, P52 Comparative Studies of Economies

Introduction

The view that institutional rigidities in labour markets are at the root of Europe's unemployment problems has become the mainstream view in economics and the public policy discourse. Originally espoused by the OECD Jobs Study (1994), this view has found support in a large body of literature (Scarpetta, 1996; Nickell, 1997, 2005; Siebert, 1997; Elmeskov *et al.*, 1998; Nunziata, 2002; IMF, 2003; Bélot and van Ours, 2004). While the OECD has subsequently toned down its initial recommendations about across-the-board institutional deregulation, their empirical studies continue to emphasize a link between certain labour market rigidities and unemployment (Bassanini and Duval, 2006; OECD, 2006). Although studies that support this view differ in their conclusions about which labour market institutions have more of an effect, they generally agree that deregulation is needed to fight unemployment. Despite its popularity, however, the evidence supporting the deregulatory view is not conclusive. Some recent contributions challenge the empirical findings that inform this position and argue that there is little or no convincing evidence that links institutions to unemployment (Baker *et al*., 2005; Baccaro and Rei, 2007, Howell, 2005; Vergeer and Kleinknecht, 2012).

This paper re-assesses this debate by bringing in new evidence from a larger group of countries, which in addition to OECD economies includes the 10 new European Union (EU) members from Central and Eastern Europe (CEE).[1] The inclusion of CEE countries promises new insights for at least three reasons. First, these countries have experienced substantial institutional reforms over the past two decades, which allows us to establish more clearly the impact of institutional changes on unemployment. Second, analysing the role of labour market institutions on a larger sample increases data variability and helps to disentangle the effects of different institutional settings on unemployment. Finally, most CEE countries have been under a strong policy influence of international organizations promoting the deregulatory view. Yet, due to data limitations, the empirical evidence supporting this policy advice has been far less compelling than commonly believed. Indeed, only a few studies examine the link between institutions and unemployment in CEE, but due to the lack of data they often rely on simple correlations and cross-sectional analysis (see Cazes and Nesporova, 2003; Behar, 2009). More recent studies employing panel data techniques focus either on a sample of OECD economies that includes only a couple of CEE countries since the mid- to late-1990s (Ederveen and Thissen, 2007; Fialova and Schneider, 2009), or alternatively consider longer panels but focus exclusively on post-communist countries (Schiff *et al*., 2006; Lehman and Muravyev, 2009). Given the drawbacks associated with data availability, it is not surprising that the findings of this literature are not conclusive. While several studies find a positive link between unemployment and institutions, such as the tax wedge, employment protection and unemployment benefits, the evidence supporting these findings is rather weak. This paper addresses this lacuna in prior research. Relying on a newly constructed dataset of labour market institutions in CEE and the standard data sources for OECD economies, we re-asses the conventional view that institutional rigidities are responsible for

[1]CEE countries include Bulgaria, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia and Slovenia.

high unemployment.[2] Our objective is three-fold. First, we estimate a standard dynamic model of unemployment on data covering most EU and OECD countries since 1980. We run a battery of sensitivity checks to assess the robustness of these estimates. Second, we examine whether the determinants of unemployment are different in advanced and new market economies. Finally, we assess the possibility that institutions affect unemployment not only directly, but also through their interactions with macroeconomic shocks.

Overall, we find no compelling support for systematic deregulation. Only a few institutions, namely union density and to a lesser extent unemployment benefits, seem to be associated with high unemployment, but we show that these findings are highly fragile. Sensitivity checks suggest that the adverse effects of these institutions diminish or disappear entirely with small changes in the sample and data, or the use of alternative estimators. While we do find some indications that simultaneous reforms of unemployment benefits and the tax wedge may be beneficial, this evidence is also not immune to small changes in the sample. Meanwhile, the effects of macroeconomic controls, such as GDP growth, and the unemployment-reducing effects of wage bargaining coordination remain fairly robust. When we consider the two groups of countries separately, it becomes clear that the impact of institutions in the new market economies is even weaker: Here, most labour market institutions are already fairly liberal, and unemployment seems to be related primarily to macroeconomic developments. Evidence that institutions affect unemployment indirectly, by exacerbating the effects of macroeconomic shocks, is similarly inconclusive and highly contingent on the statistical properties of the models. In sum, our analysis demonstrates rather thin support for the deregulatory view. A simple labour market rigidity story appears too simplistic to account for developments in unemployment in the EU and OECD countries over the last three decades.

The paper is divided in seven sections. Section 1 provides an overview of unemployment trends and compares the experience of the advanced OECD economies and the new market economies from CEE. Section 2 presents the main models and hypotheses. Section 3 offers a brief summary of the data, while

[2]Some authors argue that the employment rate is a better indicator of the overall health of the labour market because joblessness may be masked by low labour force participation, different types of active labour market policies, or widespread use of early retirement options (Kenworthy, 2008, p. 62). We focus on the unemployment rate primarily because the literature that supports the deregulatory view predominantly uses this measure as an indicator of the current labour market performance. In addition, the unemployment rate arguably carries more political weight: it seems reasonable to assume that incumbent politicians are punished more by high unemployment than low employment. Finally, unemployment and employment rates are highly correlated. Given this, it is not surprising that using the employment rate as the dependent variable does not generate markedly different results from those presented here.

Section 4 outlines our estimation strategy. Section 5 provides a detailed discussion of the results and employs several sensitivity checks to assess their robustness. Section 6 discusses the possible reasons for the weak effects of institutions and for the differences in results within this literature. The last section concludes by summarizing the key findings.

Labour market institutions and unemployment in the EU and OECD countries

Figure 1 plots the evolution of unemployment and key labour market institutions in continental Western Europe, the USA, and CEE during 1980–2007 (for summary statistics, see online appendix). The first panel shows that since the early 1980s unemployment in continental Europe has been generally higher than in the USA with a particularly large gap during the 1990s. Due to the transformational recession, CEE countries experienced a dramatic jump in unemployment during the early 1990s that in many cases surpassed unemployment levels in the advanced countries. The gap even widened during the late 1990s when a second round of restructuring in CEE triggered a new increase in unemployment. More recently, however, the differences between countries have narrowed substantially. The remaining panels reveal some striking differences in the development of the main labour market institutions. While union density has been generally declining, this decline has been particularly striking in CEE where the end of compulsory union membership led to large membership losses. CEE countries also experienced a dramatic decline in the unemployment benefit replacement ratio. As unemployment soared in the early transition, CEE governments adopted a series of reforms reducing the generosity of benefits. By the late 2000s, the average replacement rate was significantly lower than in continental Europe, and very similar to the USA. Employment protection is also more liberal in CEE than in continental Europe, but not as liberal as in the USA. The gap between continental Europe and CEE has narrowed in the 2000s when the CEE countries were required to adopt a number of EU directives that define the procedures regarding collective dismissals and increased protection of temporary employees. CEE also occupies the middle position in terms of wage bargaining coordination, with a peak in the early- to mid-1990s and the subsequent decline to a level that is considerably below continental Europe. The evolution of the tax wedge displays a different trend from other variables: here CEE has experienced a steady increase so that in the late 1990s the average tax wedge exceeded the level in continental Europe and remained at a relatively high level thereafter.

In sum, this figure shows no clear indication that unemployment is a consequence of labour market rigidities. Although union density and the benefit replacement ratio in CEE show a steep decline throughout the period, unemployment has

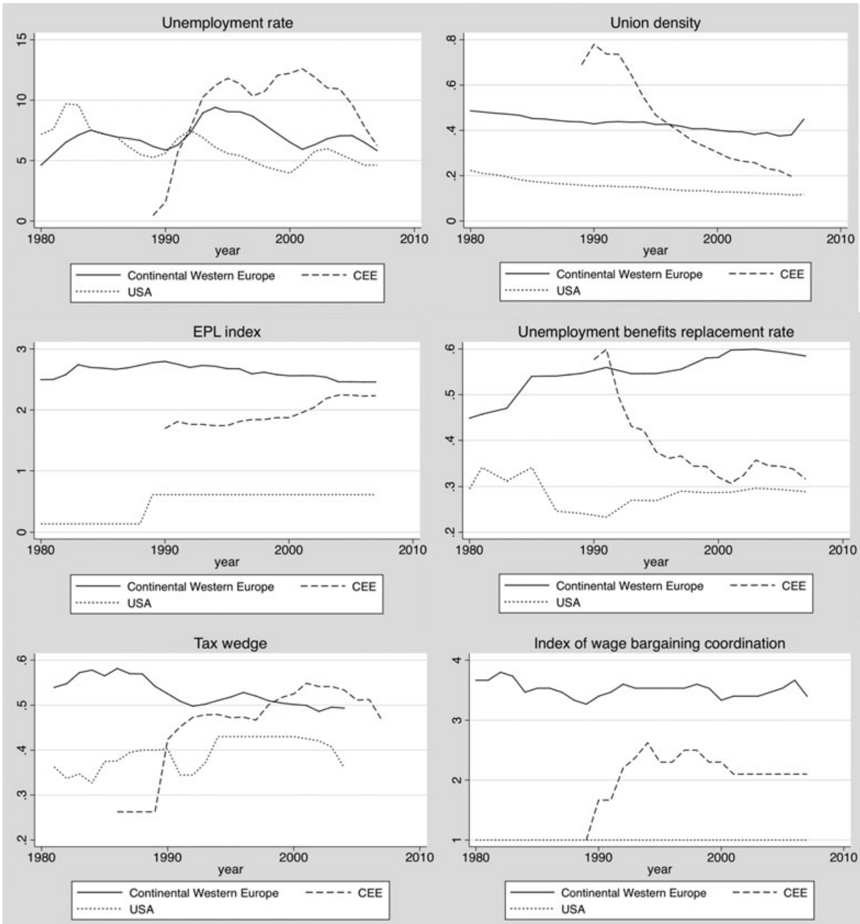

Figure 1 The evolution of unemployment and labour market institutions.
Source: See section on data below.

been increasing, with minor dips, pretty much until the mid-2000s. Similarly, unemployment in CEE was declining at the time when employment protection was strengthened and wage coordination was largely stable. The lack of a clear association between institutions and unemployment is also evident in continental Europe and the USA. The steady decline in unionization and little movement with respect to employment protection and wage bargaining coordination do not seem to offer a convincing explanation of unemployment. Similarly, the decline in unemployment during the 1990s is at odds with the moderate increases in benefit generosity. Only the tax wedge shows occasionally some association with unemployment, but apart from the USA, this association appears rather weak.

In a series of separate plots (not shown here), we examined in more detail bivariate associations between unemployment and labour market institutions in different groups of countries. Among these, only employment protection legislation (EPL) in continental Europe shows a mildly positive relationship with unemployment (the correlation coefficient of 0.14 for regular contracts and 0.36 for temporary contracts). However, a closer look reveals that this relationship is driven by only a few countries, most notably the Netherlands, and to a lesser extent Spain, Greece and Switzerland. In most countries there is no clear relationship between EPL and unemployment, and in some countries, such as Sweden, Portugal, Belgium and Germany, there is even a significant negative relationship. In addition, in countries that display a positive relationship between EPL and unemployment a causal story underlying this relationship is not quite clear. For example, while changes in EPL may have contributed to the rise of part-time employment in the Netherlands, most scholars tend to agree that the key ingredient of the 'Dutch employment miracle' was not these changes, but the willingness of unions to accept wage moderation (Visser and Hemerijck, 1997). In Spain, a moderate decline in unemployment during 1994–2007 has been often attributed to the increasing use of temporary contracts. However, recent research shows that a large gap in EPL strictness between regular and temporary contracts may actually contribute to rising unemployment by increasing the number of separations (Bentolila *et al*., 2010).

Clearly, Figure 1 provides only a very crude picture of the links between institutions and unemployment and it conceals considerable differences that exist within the groups. Within continental Europe, unemployment in the 'Big Four'—Spain, France, Italy and to a lesser extent Germany—was hovering around 10% for most of the period under consideration, and occasionally even around 20% in Spain. Greece and Belgium also struggled with unemployment for most of this period, while Finland and Denmark experienced episodes of high unemployment at different points in time. Other countries, however, had less of a problem with unemployment, and some of them (Austria, Norway and the Netherlands) even performed better than the USA during the late 1990s and early 2000s. Differences in unemployment emerged also within CEE countries, despite the fact that they embarked upon transition with generally similar labour market conditions. While Poland, Bulgaria and Slovakia had problems with high unemployment for most of the period, and the Baltic countries were affected from the mid-1990s, the Czech Republic and Slovenia managed to keep unemployment at relatively low levels. As documented by previous research, there are also notable cross-country differences in labour market institutions within these groups (Baker *et al*., 2005; Howell, 2005; Nickell *et al*., 2005; Schiff *et al*., 2006), and they do not allow straightforward conclusions about the link between institutions and unemployment. To examine this variation across countries and over time, in the remainder of this paper we turn to time-series cross-section analysis.

Models and hypotheses

We estimate a dynamic model of unemployment that has been used widely in the literature (Nickell *et al.*, 2005; see also Layard *et al.*, 1991; Nunziata, 2002; IMF, 2003; Amable *et al.*, 2006; Baccaro and Rei, 2007). In this model the unemployment rate depends on a set of labour market institutions and macroeconomic controls. The former determine the equilibrium level of unemployment, while the latter account for short-term deviations from the equilibrium level. The model has the following form:

$$u_{i,t} = \beta_0 + \beta_1 u_{i,t-1} + \Sigma_j \gamma_j x_{j,it} + \Sigma_k \eta_k z_{k,it} + \Sigma_n \delta_n v_{n,it} + \alpha_i + \lambda_t + \varepsilon_{i,t}, \quad (1)$$

where $u_{i,t}$ represents the unemployment rate in country i at time t, $u_{i,t-1}$ is the lagged unemployment rate, $x_{j,it}$ are j institutional variables, $z_{k,it}$ represent k macroeconomic controls, $v_{n,it}$ are n interactions between labour market institutions and $\varepsilon_{i,t}$ is the stochastic residual. The model also includes country dummies, α_i, which account for unmeasurable time-invariant country-specific characteristics that may influence unemployment, and year dummies, λ_t, which denote time-varying shocks affecting all countries. The lagged-dependent variable is included among the predictors to capture the persistence of unemployment and hysteresis effects (Nickell *et al.*, 2005).

The vector of institutional variables includes as follows:

$$\Sigma_j \gamma_j x_{j,it} = \gamma_1 \mathrm{EP}_{i,t} + \gamma_2 \mathrm{BRR}_{i,t} + \gamma_3 \mathrm{TW}_{i,t} + \gamma_4 \mathrm{UD}_{i,t} + \gamma_5 \mathrm{BC}_{i,t}, \quad (2)$$

where $\mathrm{EP}_{i,t}$ is employment protection legislation, $\mathrm{BRR}_{i,t}$ is the unemployment benefit replacement rate, $\mathrm{TW}_{i,t}$ is the tax wedge, $\mathrm{UD}_{i,t}$ is union density and $\mathrm{BC}_{i,t}$ is wage bargaining coordination. In the standard competitive model, employment protection legislation dampens job creation because employers are reluctant to hire new workers for fear of not being able to fire them easily when the need arises (Addisson and Texeira, 2003). However, strict EPL also increases job retention as employers make fewer layoffs during downturns. In addition, strong job protection may encourage investments in training and enhance overall productivity performance (Estevez-Abe *et al.*, 2001). Thus, the overall effect of EPL on unemployment, as Bertola (1992) has argued, is theoretically ambiguous, and it may depend on issues such as the functional form of labour demand functions, the discount rate, labour turnover and wage flexibility. The impact of unemployment benefits is generally less ambiguous. Generous benefits are thought to increase unemployment because they indicate a high reservation wage, which makes unemployed individuals both more reluctant to seek actively for jobs and to accept available jobs (Nickell, 1997; Holmlund, 1998). In addition, generous benefits may contribute to unemployment by making unions more resolute in pushing for higher wages

(Layard *et al.*, 1991).[3] The tax wedge, the difference between the labour cost to employers and the take-home wage for employees, is generally expected to influence labour market performance negatively by reducing the demand for labour. However, theoretically the distribution of taxes between employers and labour determines the actual impact of the tax wedge (Nickell, 1997). If employees carry most of the tax burden, this variable alone is not likely to reduce labour demand. At the same time, the impact on labour supply is indeterminate since a low take-home pay may either reduce workers' incentive to accept jobs and keep the existing ones, or it may motivate them to seek additional jobs. Union density indicates union bargaining power. In the orthodox view, unions tend to raise wages, and therefore a high share of workers belonging to unions is expected to increase unemployment, particularly in contexts of a highly elastic labour supply. Strong unions are also associated with compressed wage structures, which may reduce the prospects for employment of low-skill workers (Rueda and Pontusson, 2000). In contrast, the effect of wage bargaining coordination is generally considered to be beneficial for labour market performance. Because unions in coordinated systems internalize the externalities of their wage policies, it is expected that real wages, and thus unemployment, will be lower than in systems characterized by uncoordinated bargaining (Hall and Franzese, 1998).

The macroeconomic controls include as follows:

$$\Sigma_k \eta_k z_{k,it} = \eta_1 \mathrm{CPI}_{i,t} + \eta_2 \mathrm{GDP}_{i,t} + \eta_3 \mathrm{TOT}_{i,t} + \eta_4 \mathrm{RIR}_{i,t}, \tag{3}$$

where CPI is the change in inflation, GDP is GDP growth, TOT is the terms of trade and RIR is the real interest rate. Change in inflation captures the influence of economic cycles (Nickell, 1997). Following the logic of the Phillips curve, this variable should be negatively related to unemployment in the short run. Because there are some concerns about the suitability of this variable in the context of transition economies (Cazes and Nesporova, 2003), we include GDP growth as an additional control. A fall in output should be associated with higher unemployment values. The terms of trade variable should have a negative relationship with unemployment. A deterioration of the terms of trade requires a downward adjustment of real wages. If wages do not respond accordingly, unemployment is likely to increase. The real interest rate affects capital accumulation and can cause shifts in labour demand. This variable should be positively associated with unemployment, because an increase in real interest rates is likely to reduce aggregate demand, thereby generating higher unemployment rates (Baker *et al.*, 2005).

Finally, we include three types of interactions among institutional variables that allow us to examine possible complementarities across labour market reforms (see

[3] However, if generous benefits increase the effectiveness of the job matching process, their impact will be less clear and theoretically indeterminate.

Amable *et al*., 2006; Bassanini and Duval, 2006). The first is the interaction between the tax wedge and the generosity of unemployment benefits. Bélot and van Ours (2004) argue that simultaneous reductions in the tax burden and unemployment benefits have been important ingredients in reforms in countries that managed to reduce unemployment during the 1990s (see also Bassanini and Duval, 2006). The theoretical rationale is that if workers shoulder most of the labour taxes, the incentives of job seekers to invest heavily in job search will be lower provided that unemployment benefits are generous. The other two interactions include employment protection on the one hand, and the tax wedge and unemployment benefits on the other. These interactions help to examine whether the impact of employment protection, which is theoretically ambiguous, may be high when associated with another institutional rigidity. In theory, one channel through which these interactions work is the interdependence of the search intensities of workers and employers. High labour taxes may discourage vacancy posting because they reduce either the demand for or supply of labour. By increasing the costs of hiring and firing, strict employment protection also discourages vacancy posting. Consequently, the search intensity of workers may be reduced because the likelihood of finding a job is smaller. The adverse effects of these two institutions therefore may amplify each other. The interaction between employment protection and unemployment benefits follows a similar logic. Strict employment protection may reduce vacancy postings, and this effect may be amplified by generous benefits. The latter institution reduces workers' incentives to look for jobs, which consequently may also discourage vacancy posting (see OECD, 1999; IMF, 2003; Amable *et al*., 2006). In addition, we also estimate a number of alternative models that include additional interactions and variables. Section five outlines in detail the rationale for and the results of these models. All interactions are specified as deviations from cross-country and over-time sample means. Using such formulation, the coefficients of these institutions in levels can be interpreted as the coefficients of a country that has the average level of a given institution (Nunziata, 2002, p. 9). A negative and significant interaction coefficient between two variables that increase unemployment would suggest reform complementarity (see Bassanini and Duval, 2006, p. 21).

Finally, we supplement this analysis with an examination of interactions between institutions and macroeconomic shocks. As argued by Blanchard and Wolfers (2000, p. C17), labour market institutions may affect the impact of shocks on unemployment as well as the persistence of unemployment in response to shocks. For example, with respect to the first channel, a slowdown in productivity growth may result in unemployment unless wages are adjusted downwards, and this adjustment may be more difficult in systems with strict employment protection or generous unemployment benefits. Similarly, once the adverse shocks generate an increase in unemployment, the institutions may prolong the time needed for unemployment to return to its normal level. To examine this hypothesis, we re-estimate the Blanchard and Wolfers model. This model captures the interaction between institutions and common unobservable shocks, which are treated as time effects:

$$u_{i,t} = \lambda_t(1 + \Sigma_j \gamma_j x_{j,it}) + \alpha_i + \varepsilon_{i,t}, \tag{4}$$

where $u_{i,t}$ is unemployment in country i at time t, α_i is the country effect for country i, λ_t is the country-unvarying time effect for year t and x_j is the same set of institutions considered in the baseline linear model. The effects of common shocks depend on labour market institutions, and the coefficients γ_i capture this indirect effect of institutions on unemployment.

Data

Our data cover 32 EU and OECD countries, including all current EU member states (apart from Cyprus and Malta), Norway, Switzerland, the USA, Canada, Australia, New Zealand and Japan during 1980–2009. The series for CEE countries are shorter, starting roughly at the beginning of their democratic transitions.[4] This is a significantly larger sample than commonly used in the literature. While recent analyses by Feldmann (2009) and Bernal-Verdugo *et al.* (2012) include 73 and 97 countries, respectively, their time series are rather short. The former focuses on 3 years only, while the latter uses series that vary from 3 to 12 years.

Our dependent variable captures the number of unemployed persons as a percentage of the labour force and is based on labour force surveys (IMF World Economic Outlook and EBRD). Among our independent variables, some are newly constructed. An important contribution of our analysis is that it includes the longest and previously unavailable series that measure the strictness of EPL in CEE countries on a yearly basis during 1990–2009. This is in contrast to the series provided by the OECD, which are interpolated from a few data points. As such, these data reflect more accurately the differences in the timing and the extent of EPL reforms.[5] We combine our data for CEE countries with Allard's (2010) EPL index for advanced economies, which also captures annual changes in legislation and is based on the same methodology. The series on the unemployment benefit replacement rates for CEE countries is also newly constructed based on the scheme used by the OECD. These data capture the gross replacement rates in the first year of unemployment across two levels of earnings (67 and 100% of

[4]For reasons of comparability, the analysis excludes the first 3 years of post-communist economic transformation when these countries experienced profound macroeconomic shocks.

[5]These data can be accessed at http://store.data-archive.ac.uk/store/collaborativeCollectionEdit.jsp?collectionPID=archive%3A598&tabbedContext=allCollection.

average wage). As a check, we also re-ran all the models with the recently released data on the net replacement rate (van Vliet and Caminada, 2012). The results were not markedly different from those presented below.

In addition, the analysis uses new data on the tax wedge provided by Labartino (2010). This database provides longer and more complete series for this sample of countries than the OECD and Eurostat data. Data on union density and wage coordination are taken from Visser (2009). Data on macroeconomic controls come from the International Monetary Fund's International Financial Statistics (GDP), the World Bank's World Development Indicators (real interest rate), OECD National Accounts data files (inflation) and the European Commission's AMECO database (terms of trade). Data on central bank independence, used in one model, come from Crowe and Meade (2007). The number of countries covered in the final models was governed by data availability. The main specifications outlined above are based on a sample of 26 countries. The lower number of observations in these models is primarily a consequence of missing data on the tax wedge for Bulgaria, the Czech Republic, Slovenia, New Zealand and Switzerland, and on employment protection for Luxembourg.

Estimation strategy

We compare the results from two different estimators. The first is a panel weighted least squares estimator (PWLS), which is the most commonly used estimator in the literature that supports the deregulatory view. This model assumes country-specific heteroskedasticity and employs a Prais–Winsten transformation to address a first order (AR1) autoregressive structure in the errors (a common estimated rho). In comparison to the Parks estimator, which produces severely underestimated standard errors in analyses where T is not significantly larger than N, this feasible generalized least squares (FGLS) estimator has better properties. However, this estimator is not designed to correct for contemporaneously correlated errors, which characterize our data. In such cases, PWLS may also suffer from overoptimistic errors, which is why we prefer the ordinary least squares procedure with panel corrected standard errors (OLS-PCSE) (Beck and Katz, 1995). Used widely in comparative political economy, this estimator applies OLS with corrected standard errors to control for common properties of this type of data, including panel heteroskedasticity and contemporaneous correlation of the error terms. All models include country and year effects, as indicated by the *F*-test for their inclusion. Given that the inclusion of the lagged-dependent variable can make the fixed-effect estimator biased due to the correlation between the demeaned-lagged-dependent variable and the error term (Nickell, 1981), we also estimated the least squares dummy variable model corrected for the so-called Nickell bias, as suggested by Kiviet (1995). The results of models that use the Kiviet estimator (obtainable upon request) do not differ appreciably from the results reported below. This supports the conclusions of Beck and Katz (2011) that this bias is small in panels where T is around twenty or more. Finally, Equation (4), which captures the interaction between shocks and institutions, is estimated by non-linear least squares, as proposed by Blanchard and Wolfers (2000).

Empirical results

Table 1 reports results for four different specifications across the two estimators. Models (1) and (2) are our baseline specification. Models (3) and (4) add the interactions between the tax wedge, unemployment benefits and employment protection.

The coefficients of the lagged-dependent variable are high (0.81–0.83), indicating considerable persistence of unemployment, but also potential problems with stationarity. Unit root tests, however, suggest that most series are stationary.[6] The Augmented Dickey Fuller (204.28) and Philips-Perron (83.34) tests (Maddala and Wu, 1999) reject the null hypothesis of non-cointegration at the 1% level. The dynamic specification does not eliminate fully serial correlation in the residuals, but Monte Carlo evidence indicates that the associated bias is unlikely to be substantial (Beck and Katz, 2011).

At first glance, columns 1–4 suggest that labour market institutions have a role in explaining unemployment. At the same time, however, it is evident that support for the deregulatory view is far from strong. Among the institutions, only union density and to a lesser extent unemployment benefits are consistently positively associated with unemployment. While the finding about union density seems to hold across the estimators and specifications, the significance of the benefit replacement rate is more sporadic when FGLS is used. The tax wedge has a noticeable effect only in interactions with other institutions. Specifically, Models (3) and (4) signal reform complementarity, suggesting that a reduction in the tax wedge may augment unemployment-reducing effects of replacement rate cuts. Employment protection legislation does not have any discernable impact on unemployment, regardless of the choice of estimators and specifications. Finally, wage coordination shows unemployment-reducing properties, and this finding is robust to changes in specifications and estimators. The coefficients for all macroeconomic controls are signed as expected, but only growth is robustly significant across the models. The results are similar if we use the output gap instead of GDP growth.

[6]We ran unit root tests with one lag and two lags, with and without drift, with and without trend and with and without demean option. Central bank independence, used in one model, is the only variable for which it is clear that the null of a unit root cannot be rejected.

Table 1 Determinants of unemployment in the EU and OECD countries, 1980–2006

	(1) FGLS het (AR1)	(2) OLS-PCSE	(3) FGLS het (AR1)	(4) OLS-PCSE	(5) FGLS het (AR1)	(6) OLS-PCSE	(7) FGLS het (AR1)	(8) OLS-PCSE
L.ur	0.821***	0.813***	0.827***	0.818***	0.825***	0.825***	0.816***	0.814***
	(0.028)	(0.026)	(0.028)	(0.026)	(0.028)	(0.024)	(0.028)	(0.029)
BRR	0.481	1.306***	0.136	0.810**	0.491	1.248***	0.707	1.354**
	(0.601)	(0.428)	(0.642)	(0.405)	(0.616)	(0.432)	(0.644)	(0.576)
EP	−0.024	0.166	0.066	0.186	−0.009	0.150		
	(0.156)	(0.220)	(0.169)	(0.233)	(0.159)	(0.240)		
TW	0.821	0.981	1.150	1.582**	0.888	1.324*	0.836	1.050
	(0.661)	(0.698)	(0.699)	(0.747)	(0.668)	(0.706)	(0.669)	(0.720)
BC	−0.146***	−0.236***	−0.117**	−0.206***	−0.111*	−0.155**	−0.162***	−0.254***
	(0.055)	(0.068)	(0.057)	(0.073)	(0.064)	(0.073)	(0.056)	(0.064)
UD	4.393***	2.855***	4.305***	3.542***	4.417***	2.470***	4.288***	3.014***
	(1.330)	(0.755)	(1.312)	(0.713)	(1.368)	(0.574)	(1.364)	(0.781)
GDP	−0.262***	−0.284***	−0.259***	−0.287***	−0.264***	−0.285***	−0.247***	−0.279***
	(0.022)	(0.019)	(0.022)	(0.019)	(0.023)	(0.018)	(0.023)	(0.019)
CPI	−0.012	−0.003	−0.019	−0.009	−0.012	0.001	−0.013	−0.002
	(0.022)	(0.036)	(0.022)	(0.038)	(0.022)	(0.035)	(0.022)	(0.036)
TOT	−0.003	−0.019***	−0.002	−0.016**	−0.004	−0.020***	−0.003	−0.019**
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)	(0.007)	(0.008)
RIR	0.018	0.028	0.018	0.029	0.021	0.032*	0.019	0.028
	(0.014)	(0.020)	(0.014)	(0.019)	(0.014)	(0.020)	(0.014)	(0.020)
EP*BRR			0.176	0.562				
			(0.532)	(0.523)				
EP*TW			0.557	0.091				
			(0.836)	(0.931)				
BRR*TW			−12.873***	−12.228***				
			(3.990)	(4.037)				

Continued

Table 1 Continued

	(1) FGLS het (AR1)	(2) OLS-PCSE	(3) FGLS het (AR1)	(4) OLS-PCSE	(5) FGLS het (AR1)	(6) OLS-PCSE	(7) FGLS het (AR1)	(8) OLS-PCSE
CBI					0.025	−0.604*		
					(0.290)	(0.343)		
BC*CBI					0.187	0.554***		
					(0.203)	(0.151)		
EPR							0.023	0.099
							(0.157)	(0.118)
EPT							0.024	0.087
							(0.087)	(0.120)
EPR*EPT							−0.050	0.011
							(0.091)	(0.069)
Constant	−2.610	2.502***	−2.690	1.908**	−2.696	2.602***	−2.597	2.532***
	(3.254)	(0.850)	(3.274)	(0.822)	(3.099)	(0.878)	(3.249)	(0.821)
Country and year effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	451	451	451	451	451	451	448	448
R^2		0.949		0.949		0.950		0.949
Estimated rho	0.353		0.357		0.348		0.359	
Durbin M test for remaining serial correlation of residuals	Coeff. 0.092 *P*-value = 0.026	Coeff. 0.092 *P*-value = 0.026	Coeff. 0.155 *P*-value = 0.000	Coeff. 0.155 *P*-value = 0.000	Coeff. 0.130 *P*-value = 0.002	Coeff. 0.130 *P*-value = 0.002	Coeff. 0.113 *P*-value = 0.008	Coeff. 0.113 *P*-value = 0.008
Multicollinearity test: mean VIF	1.25	1.25	1.43	1.43	1.30	1.30	1.37	1.37

Note: BRR, unemployment benefit replacement rate; EP, employment protection; TW, tax wedge; BC, wage bargaining coordination; UD, union density; GDP, GDP growth; CPI, change in inflation; TOT, terms of trade; RIR, real interest rate; CBI, central bank independence. Standard errors in parentheses: $*P < 0.10$, $**P < 0.05$, $***P < 0.01$. Wald tests on country and time dummies, macro controls and interactions reject the null hypothesis at 1%. The Breusch–Pagan test for cross-sectional independence [$Chi^2(325) = 650$] and group-wise heteroskedasticity [$Chi^2(26) = 214.05$] reject the null. Augmented Dickey Fuller
(204.28) and Phillips Perron
(83.34) tests imply rejection of null hypothesis of cointegration at 1%.

Columns 5–8 augment the baseline model with further interactions that the previous literature found to be significant in explaining unemployment. Models (5) and (6) include the interaction between wage coordination and central bank independence. The latter variable, when observed in isolation is generally expected to increase unemployment, but it has been shown that its adverse effect tends to be lower when wage setting is highly coordinated (Hall and Franzese, 1998). In such contexts, the bargaining actors are more sensitive to the likely response of monetary policy regarding wage settlements than in non-coordinated systems. This interaction appears significant only in the OLS-PCSE model, but its positive coefficient is at odds with the standard view in the literature. However, we do not place much credence in this finding given that the significance of this coefficient is evidently sensitive to the choice of estimators as well as specifications. In models that include only inflation or GDP growth as the only control, this interaction is not significant. Models (7) and (8) examine in more detail the impact of employment protection legislation. Here, we disaggregate this variable on rules for regular and temporary contracts. In this way, we consider the possibility that the insignificant coefficient of employment protection in previous columns may mask two opposite effects, namely that EPL on regular contracts increases unemployment, while EPL on temporary contracts pushes in the opposite direction (Bassanini and Duval, 2006). We find no support for this hypothesis.

In separate regressions, we also estimated models that include the duration of unemployment benefits, which (following the FRDB Social Reforms database) reflects the number of months during which benefits are payable. We found no evidence that this variable increases unemployment either independently or through the interaction with the replacement rate. We also experimented by including the minimum wage, but this variable was not consistently significant in any models. Finally, we examined the impact on unemployment of the coverage by collective agreements. This is because union density may underestimate the strength of unions in countries where a low proportion of the labour force belongs to unions, but a large share is covered by collective agreements (e.g. France). Given the lack of full time series for bargaining coverage for all countries we could only use the average values of this variable for the whole period. Following Bassanini and Duval (2006), we created a dummy for high bargaining coverage where countries with coverage exceeding 50% were assigned score 1. Since this variable is time-invariant, its impact can be gauged only through interactions with other institutions. No interaction, however, turned out to be consistently significant across the models.

In sum, the main conclusions from Table 1 remain unaffected in the alternative specifications. Among institutions union density remains consistently associated with unemployment, while wage coordination helps to reduce unemployment. The results also suggest that generous benefits may play a role, but this finding is less stable and specification-dependent. It needs to be noted though that although union density and benefit generosity seem to be associated with higher unemployment, the magnitude of these effects is relatively small. The estimates of our baseline OLS-PCSE model imply that the impact of a 20 percentage point increase (equal to one standard deviation) in union density yields an increase in the unemployment rate of 0.57 percentage points. Similarly, a 19.4 percentage increase in the benefit replacement rate is associated with an increase in the unemployment rate of only 0.25 percentage points.

Before exploring in more detail the sensitivity of these results, the potential endogeneity of institutions needs to be addressed. Although the mainstream literature focuses on the impact of institutions on unemployment, it is reasonable to think that the state of the labour market and employment prospects influence decisions of policy makers about reforms. For example, Howell and Rehm (2009) have shown that the causality may run from unemployment to benefit generosity rather than vice versa. Although this hypothesis is plausible, Granger causality tests on our data show no clear evidence of reverse causality.[7] One could, however, argue that endogeneity problems may be still present if omitted variables influence simultaneously institutions and unemployment. In separate regressions, we checked this possibility by re-estimating our baseline model using a difference GMM estimator (Arellano and Bond, 1991), where institutions are instrumented with their lagged values. The results are not very different from our baseline point estimates.[8]

Sensitivity analysis

The finding that union density and benefit replacement rates are positively associated with unemployment is in line with the mainstream literature on the subject. But how robust are these findings? As we saw, the results are generally not sensitive to changes in specifications. In addition, the inclusion of fixed effects is supposed to capture possible country- and year-specific influences.

[7] We performed Granger causality tests by estimating models with two lags of the unemployment rate and labour market institutions. Benefit generosity and union density are of particular interest, because their baseline point coefficients are positive and significant. The *F*-statistic of the two lagged term of the explanatory variables is not significant, implying that causation does not run from unemployment to institutions. Results obtainable upon request.

[8] Because this estimator is designed for small T panels, we follow the approach used by Bassanini and Duval (2009) and estimate these models on 5-year averaged data. Levels of endogenous variables dated $t-2$ and earlier are used as instruments. In the first model, all institutions are treated as endogenous variables, while in the second model only those institutions that have a significant impact on unemployment in the baseline models are treated as endogenous. The results of these models are very similar, although the first one is more fragile due to a higher number of instruments. In both models benefit generosity and wage bargaining coordination retain their significance.

Nonetheless, given the relatively small sample size, it is still possible that some countries or even individual observations greatly influence the coefficient estimates. Thus, we perform two additional checks. First, we run a jackknife analysis on the models presented in Table 1. This procedure re-estimates repeatedly the same models by omitting from the sample one country each time. Second, we re-estimate the same models by removing a small number of potentially influential observations that were identified through a combination of a visual inspection and dffits and covratio cut-offs. This exercise revealed six observations that may have a disproportional influence on the results: Lithuania in 1995 and 2002, Spain in 2001, Latvia in 1996, Romania in 2004 and Finland in 1993. Table 2 shows the outcomes of these checks for the baseline model. Column 3 reports the original estimates. Column 6 shows the results when the six outliers are excluded from the sample. The remaining columns report the maximum and minimum value of coefficients obtained by jackknifing, and the country that was omitted when those coefficients were obtained.

As evident, the sensitivity analysis implies significant differences in the substantive conclusions about the impact of institutions. Although none of the originally significant coefficients changes the sign in the jackknife analysis, the results are clearly not robust as the key coefficients become insignificant upon exclusion of a single country. In particular, the results regarding union density, which showed up as the main culprit of unemployment, are fragile and hinge entirely on the presence of one country in the sample—omitting Lithuania reduces the coefficient considerably and makes it statistically indistinguishable from zero. Similarly, the benefit replacement ratio also becomes insignificant if Lithuania, Portugal or Austria is excluded from the sample. Among initially significant institutions only wage bargaining coordination remains robust, but this institution is associated with lower unemployment. Turning to macroeconomic controls, GDP growth retains its negative sign and significance irrespective of the changes in the sample. The main findings of the jackknife analysis are corroborated when we exclude the six country-years identified as outliers. As column 6 shows, labour market institutions do not seem to have a direct detrimental effect. Again bargaining coordination remains robustly associated with lower unemployment.

The results (not shown) of the sensitivity analysis for the main interaction model (Table 1, column 4) are less conclusive. While the jackknife analysis shows that the coefficient of the interaction between the tax wedge and the benefit replacement becomes insignificant when Spain is omitted, this coefficient retains its significance when the six outliers are removed from the sample. In sum, contrary to the initial findings, the sensitivity analysis indicates that individual institutions do not have a clear detrimental effect on unemployment. At the same time, evidence about reform complementarity between the tax wedge and unemployment benefits reforms remains inconclusive.

Table 2 Jackknife analysis and the exclusion of outliers for the baseline model

	Minimum	Country	Estimate	Maximum	Country	Estimate w/o outliers
L.ur	0.786*** (0.028)	SP	0.813*** (0.026)	0.910*** *0.029)	LT	0.904*** (0.032)
BRR	0.169 (0.388)	LT (PT, AUT)[a]	1.307*** (0.428)	2.538*** (0.347)	IT	0.321 (0.401)
EP	−0.160 (0.246)	SP	0.166 (0.220)	0.385 (0.238)	SWE	0.110 (0.239)
TW	−0.086 (0.663)	LT	0.981 (0.698)	2.444*** (0.793)	SP (GR, IT)[b]	0.319 (0.723)
BC	−0.280*** (0.076)	AUS	−0.236*** (0.068)	−0.177*** (0.068)	LT	−0.144** (0.071)
UD	0.437 (0.861)	LT	2.855*** (0.755)	4.889*** (0.878)	SP	0.774 (0.878)
CPI	−0.032 (0.030)	IRL	−0.003 (0.036)	0.019 (0.0423)	CAN	−0.290*** (0.017)
GDP	−0.341*** (0.019)	RO	−0.284*** (0.019)	−0.262*** (0.018)	PL	−0.003 (0.034)
TOT	−0.024*** (0.007)	NO	−0.019*** (0.007)	−0.007 (0.006)	LT (SP, SWE)[c]	−0.008* (0.005)
RIR	0.002 (0.024)	RO	0.028 (0.020)	0.061*** (0.019)	LT (SK)[d]	0.056*** (0.017)

Notes: Entries in the first five columns are OLS-PCSE coefficient estimates from the baseline model (Table 1, column 2), together with minimum and maximum coefficient estimates obtained by re-estimating the model so that each country is omitted one at a time. The last column presents the estimates of the model where six outliers (Lithuania 1995 and 2002, Spain 2001, Latvia 1996, Romania 2004 and Finland 1993) are excluded from the sample.

[a]Exclusion of Portugal or Austria also makes the coefficient of the replacement rate insignificant.

[b]Apart from Spain, the tax wedge becomes significant when Greece or Italy are dropped from the sample, albeit with a smaller coefficient (1.346 and 1.597, respectively).

[c]The terms of trade coefficient becomes also insignificant with the exclusion of either Spain or Portugal.

[d]The coefficient of the real interest rate turns marginally significant when Slovakia is omitted (0.026).

Do the determinants of unemployment differ between advanced and new market economies?

The previous section has shown that the estimates are highly sensitive to changes in the sample. The overall results, therefore, can conceal potentially different effects of institutions in different groups of countries. In this section, we re-assess this finding by re-estimating the main models for the advanced and new market economies separately. This additional check is also warranted on substantive grounds. Although we control for the initial transition shock in CEE countries by excluding the first three years of the transition from the analysis, there may be still good reasons to believe that the key culprits of unemployment are different in the two regions. The literature suggests that trade unions in CEE are weaker than in continental Europe and that wage bargaining is largely uncoordinated (Cazes and Nesporova, 2003). The welfare states, at the same time, do not appear very generous and support for the unemployed has been generally much more modest (Schiff *et al*., 2006). Our data also indicate that on the whole employment protection is fairly liberal in CEE. Thus, apart from the tax wedge, which remains high, this general evidence suggests that CEE countries do not seem to suffer from excessive labour market rigidities. Given that unemployment in CEE has been generally higher and labour market rigidities less pronounced, the effects of institutions on unemployment should be even weaker than in the advanced economies.

Table 3 reports the estimates of the fully dummy-interactive model, which allows us to see the results of the main models from Table 1 for the advanced and new market economies separately. The first four columns show the results for the baseline model and the main interaction model across the two estimators. Columns 5–8 repeat this exercise for the sample that excludes the six outliers identified earlier. The results suggest that labour market institutions appear to be more influential in the advanced economies. Union density and to a lesser extent benefit replacement ratios are both consistently associated with high unemployment. At the same time, wage bargaining coordination remains negative and significant in most models. As in all previous models, GDP growth remains robustly and negatively associated with unemployment. The interaction between the tax wedge and the benefit replacement rate suggests the potential for reform complementarity, but this result is evidently no longer so robust when we exclude the outliers.

The link between the labour market institutions and unemployment is less evident in CEE. The coefficients on the interactions with the CEE dummy capture the difference between the two groups of countries. In other words, the coefficient of a given variable for CEE is equal to the sum of its interaction with the CEE dummy and the corresponding coefficient for the advanced economies. As evident, the terms of trade is the only variable that seems to have a clear impact on unemployment. The tax wedge seems to have influence only in the models

Table 3 Differences between the advanced and new market economies

	(1) FGLS het (AR1)	(2) OLS-PCSE	(3) FGLS het (AR1)	(4) OLS-PCSE	(5) FGLS het (AR1)	(6) OLS-PCSE	(7) FGLS het (AR1)	(8) OLS-PCSE
L.ur	0.875***	0.917***	0.888***	0.928***	0.865***	0.900***	0.874***	0.910***
	(0.025)	(0.018)	(0.025)	(0.019)	(0.024)	(0.022)	(0.025)	(0.025)
BRR	1.296**	1.279***	1.428*	0.809**	1.216**	1.215***	1.443**	0.911*
	(0.606)	(0.300)	(0.733)	(0.387)	(0.586)	(0.325)	(0.708)	(0.484)
EP	−0.050	0.260	0.047	0.292	−0.097	0.179	−0.021	0.224
	(0.138)	(0.181)	(0.148)	(0.187)	(0.133)	(0.209)	(0.144)	(0.213)
TW	−0.003	−1.391***	0.111	−0.590	0.170	−0.913	0.312	−0.203
	(0.571)	(0.509)	(0.602)	(0.484)	(0.539)	(0.579)	(0.580)	(0.531)
BC	−0.149***	−0.155**	−0.120**	−0.112	−0.141***	−0.143**	−0.113**	−0.098
	(0.050)	(0.065)	(0.052)	(0.071)	(0.049)	(0.068)	(0.051)	(0.073)
UD	4.455***	1.796***	3.599***	1.736***	4.467***	2.232***	3.838***	1.998***
	(1.282)	(0.461)	(1.260)	(0.448)	(1.228)	(0.548)	(1.232)	(0.498)
GDP	−0.293***	−0.321***	−0.292***	−0.320***	−0.294***	−0.309***	−0.290***	−0.307***
	(0.023)	(0.019)	(0.023)	(0.019)	(0.022)	(0.021)	(0.022)	(0.021)
CPI	−0.009	−0.003	−0.012	−0.010	−0.009	−0.004	−0.011	−0.012
	(0.022)	(0.029)	(0.022)	(0.029)	(0.022)	(0.028)	(0.022)	(0.029)
TOT	−0.003	−0.003	−0.001	0.001	−0.003	−0.003	−0.002	0.001
	(0.007)	(0.004)	(0.007)	(0.005)	(0.006)	(0.004)	(0.006)	(0.005)
RIR	0.021	0.052**	0.026	0.061***	0.022	0.049**	0.025	0.056**
	(0.016)	(0.022)	(0.016)	(0.022)	(0.015)	(0.022)	(0.016)	(0.022)
L.ur*CEE	−0.435***	−0.470***	−0.434***	−0.490***	−0.073	−0.132	−0.120	−0.211**
	(0.093)	(0.064)	(0.095)	(0.050)	(0.091)	(0.092)	(0.089)	(0.103)
BRR*CEE	2.094	0.780	6.969**	7.673***	−0.836	−1.142	4.265	3.777
	(2.229)	(1.504)	(3.148)	(1.907)	(1.830)	(1.578)	(3.014)	(2.605)

EP*CEE	−1.163	−1.569	−0.885	−1.403	−0.639	−0.742	−0.170	−1.539
	(1.017)	(1.239)	(1.384)	(0.883)	(0.627)	(1.051)	(0.790)	(1.082)
TW*CEE	7.526	10.198*	6.717	8.907*	7.036*	8.599*	9.310**	11.211***
	(5.545)	(6.124)	(5.470)	(4.908)	(4.168)	(5.053)	(4.367)	(3.765)
BC*CEE	−0.565	−0.738	0.089	0.144	0.176	0.174	0.737	0.829
	(0.692)	(0.923)	(0.690)	(0.823)	(0.523)	(0.731)	(0.566)	(0.695)
UD*CEE	2.108	4.928	−3.042	−1.694	−4.640	−0.758	−5.657	−4.412
	(6.603)	(8.268)	(7.070)	(6.783)	(5.180)	(6.021)	(5.345)	(5.348)
GDP*CEE	0.043	0.011	0.117	0.085*	0.109*	0.124***	0.141**	0.144***
	(0.085)	(0.047)	(0.084)	(0.045)	(0.061)	(0.042)	(0.066)	(0.056)
CPI*CEE	−0.070	0.001	−0.062	0.012	0.189	0.179	0.247	0.360
	(0.252)	(0.348)	(0.256)	(0.252)	(0.213)	(0.396)	(0.229)	(0.297)
TOT*CEE	−0.150***	−0.167***	−0.149***	−0.172***	−0.101**	−0.122***	−0.143***	−0.159***
	(0.056)	(0.051)	(0.054)	(0.031)	(0.041)	(0.041)	(0.043)	(0.038)
RIR*CEE	−0.005	−0.015	−0.032	−0.057*	−0.019	−0.025	−0.040	−0.057*
	(0.036)	(0.046)	(0.037)	(0.029)	(0.027)	(0.039)	(0.029)	(0.031)
CEE	24.048***	—, —	21.120***	—, —	12.126**	—, —	10.878*	—, —
	(7.578)		(7.374)		(5.887)		(5.860)	
EP*BRR			−0.407	0.262			−0.453	0.103
			(0.525)	(0.461)			(0.510)	(0.530)
EP*TW			−0.220	−0.428			−0.304	−0.552
			(0.670)	(0.738)			(0.637)	(0.737)
BRR*TW			−6.214*	−10.564***			−4.998	−9.425***
			(3.558)	(3.585)			(3.463)	(3.475)
EP*BRR*CEE			14.420*	14.108***			10.818*	4.402
			(8.100)	(3.712)			(5.814)	(5.437)
EP*TW*CEE			21.158	31.627**			17.850*	28.791**
			(13.232)	(13.618)			(10.169)	(13.070)
BRR*TW*CEE			68.352	70.027**			24.619	30.800
			(46.596)	(35.626)			(35.642)	(36.865)

Continued

Table 3 Continued

	(1) FGLS het (AR1)	(2) OLS-PCSE	(3) FGLS het (AR1)	(4) OLS-PCSE	(5) FGLS het (AR1)	(6) OLS-PCSE	(7) FGLS het (AR1)	(8) OLS-PCSE
Constant	−2.995	1.352***	−4.937**	0.876*	−2.501*	11.281*	−4.327***	10.684**
	(2.005)	(0.505)	(2.130)	(0.525)	(1.378)	(6.833)	(1.542)	(4.481)
Country and year effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
No. of observations	451	451	451	451	440	440	440	440
No. of countries	26	26	26	26	26	26	26	26
R^2		0.965		0.969		0.974		0.976

Notes: Estimates in columns 1–4 are based on the full sample. Columns 5–8 exclude the outliers. Standard errors in parentheses:* significant at 10%, ** significant at 5%, *** significant at 1%.

that exclude the six outliers. Union density and bargaining coordination, which matter for unemployment in the advanced economies, have no discernable impact in CEE and are even 'wrongly' signed in most models. While the interactions between employment protection on the one hand, and benefit replacement rates and the tax wedge on the other appear as statistically significant in several models, neither of these interactions is robust to changes in estimators or the exclusion of outliers. On the whole, these findings suggest that institutions in CEE play even less of a direct role in explaining aggregate unemployment than in the advanced economies.

As above, we also assessed the robustness of these results through a jackknife analysis, which showed that the results are fragile to the exclusion of individual countries. In the case of advanced economies, the benefit replacement ratio is no longer significantly associated with unemployment if either Portugal or Austria is omitted. The only institution that remains robustly associated with higher unemployment is union density, though the magnitude of the coefficient and the significance level become markedly smaller when the UK or Finland is dropped from the sample. This finding concurs with Baccaro and Rei (2007) who demonstrate that union density is the only institutional variable that displays a robust positive association with unemployment in OECD countries. Apart from union density, only GDP growth survives the jackknife procedure. Finally, the significance of the interaction between the benefit replacement rate and the tax wedge turns out to depend entirely on the presence of Spain. Only when the interactions are added one at a time does this interaction survive the jackknife procedure, though the exclusion of Spain reduces the coefficient considerably (from 10.82 to 5.05), making it significant only at the 10% level.

The results for CEE are even more fragile. The tax wedge is no longer significant if any of the following countries is excluded from the analysis: Latvia, Lithuania, Poland, Romania, Italy, Norway, Spain or the UK. The interaction between employment protection and the benefit replacement rate becomes insignificant when either Estonia or Latvia is omitted, confirming the results of the analysis that excludes the outliers. When these interactions are added to the baseline model one at a time, only the interaction between the tax wedge and employment protection survives the jackknife procedure, but this interaction becomes insignificant as we add other interaction terms. The results for other variables are also sensitive to the exclusion of individual countries, and some coefficients even change their sign depending on the sample.

Overall, the analysis shows that the results of these models are rather fragile. Most estimates suggesting an adverse impact of institutions no longer hold when we exclude particular countries or even a small number of potentially influential observations.

Interactions between institutions and macroeconomic shocks

Although we did not find strong evidence that unemployment is a direct consequence of rigid labour market institutions, it is possible that institutions play a more indirect role by amplifying the effects of economic shocks. Table 4 reports the results of the model that examines this possibility. Following Blanchard and Wolfers (2000), this model is estimated via non-linear least squares. Positive coefficients indicate that institutions exacerbate the effects of shocks, while the negative coefficient suggests that institutions mitigate the adverse effects of shocks on unemployment.

Institutions in Blanchard and Wolfers's analysis are expressed as deviations from the sample means. Following Baccaro and Rei (2007), we extend this formulation and consider annual data in both levels and deviations. In addition, because the assumption of i.i.d. residuals is unrealistic (cf. Blanchard and Wolfers, 2000, p. 20), we rely on the results obtained using Rogers robust standard errors. These correspond to White standard errors adjusted to account for the possible correlation within a cluster (i.e. country) and country-specific heteroskedasticity.[9] When data are in levels the coefficients on the time dummies (not reported) indicate the impact of shocks on unemployment in a country in which all institutional variables are set to zero. In this case, the coefficients of the institutional variables shown in columns 1 and 2 capture the additional effect of shocks on unemployment when a given institution increases by one unit. When data are in deviations the coefficients of the time dummies capture the impact of shocks in a country where all institutions are at the sample mean, and the coefficients of institutions in columns 3 and 4 capture the additional effect of shocks when the given institution increases one unit above the sample mean.

As Table 4 shows, we cannot draw strong conclusions about the indirect impact of institutions since the way in which the data are expressed and the choice of standard errors evidently has a big impact on the results. When the data are expressed in levels, institutions do not seem to amplify the effects of shocks. When the data are in deviations, the same institutions identified initially in the linear model appear significant: union density and benefit replacement ratio seem to increase the impact of adverse shocks, while bargaining coordination mitigates it. However, the estimated time effects are negative, suggesting that they make no significant contribution to the overall increase in unemployment (see Blanchard and Wolfers, 2000, p. 20).

Discussion

Despite a wide range of models and specifications, we find little support for the standard argument that unemployment is a consequence of institutional rigidities.

[9]See Petersen (2008) for the discussion of the choice of standard errors and simulation evidence that Rogers standard errors perform best with this type of data structure.

Table 4 Interactions between Shocks and Institutions: EU and OECD countries, 1980–2006

	(1) NLS levels, Rogers standard errors	(2) NLS levels, l.s standard errors	(3) NLS deviations, Rogers standard errors	(4) NLS deviations, l.s standard errors
EP				
	0.033	0.033	0.013	0.0103
	(0.340)	(0.120)	(0.102)	(0.037)
UD				
	3.829	3.829**	1.211**	1.211***
	(5.827)	(1.721)	(0.544)	(0.162)
BRR				
	1.465	1.465**	0.463*	0.463***
	(1.383)	(0.719)	(0.239)	(0.140)
TW				
	1.345	1.345	0.425	0.425**
	(2.440)	(0.947)	(0.381)	(0.179)
BC				
	−0.260	−0.260***	−0.082**	−0.082***
	(0.208)	(0.084)	(0.040)	(0.021)
No. of observations	500	500	500	500
R^2	0.806	0.806	0.806	0.806
Adj. R^2	0.798	0.783	0.798	0.783

Notes: Time and country dummies omitted.

As demonstrated, most positive associations between institutions and unemployment disappear with small changes in specifications or the sample. While these findings concur with recent research that questions the empirical evidence behind the deregulatory view (Baker *et al*., 2005; Baccaro and Rei, 2007; Howell and Rehm, 2009; Vergeer and Kleinknecht, 2012), they are at odds with a number of studies that report adverse effects of institutions on unemployment (e.g. OECD, 1994; Scarpetta, 1996; Siebert, 1997; Elmeskov *et al*., 1998; Blanchard and Wolfers, 2000; Nickell *et al*., 2005; Bernal-Verdugo *et al*., 2012). How can we explain such different findings? Three possible explanations are worth considering.

The first one is an obvious point that the data used here are different. Given that the choice of data (both measures of institutions and the sample considered in the analysis) inevitably has a large impact, findings of most studies in this literature are not strictly comparable. A related point is that not all studies use the same robustness checks. The fact that the results may be robust to variations in variable specification or the estimation method does not guarantee that they are robust to small changes of the sample. But this is unlikely to be the whole story. The second possible explanation is that labour market institutions simply do not have strong effects on unemployment. As indicated above, even when certain institutional variables display a statistically significant impact on unemployment, this impact is rather small in substantive terms. It should be noted that in a number of studies that support the deregulatory view, the size of the effect of particular institutions is also rather small, although this is often not explicitly acknowledged (see Baker *et al.*, 2005, pp. 101–103). This could be the case either because the positive and negative effects of institutions balance out or, as Freeman has argued, because 'bargaining settlements and regulations that are truly expensive to an economy' are effectively ruled out (2008, p. 25). Finally, an explanation that seems most convincing is that there is no universal cause and thus solution to unemployment. The same institution may have different effects in different countries or time periods (Hall, 2003, p. 383). In this line of reasoning, the impact of institutions is not straightforward and it may depend on the overall institutional configurations and interactions between labour markets and other spheres, such as social policy, skill regimes, and product market—characteristics that, given policy changes in these areas, may not be adequately captured by country dummies. More generally, this interpretation is in line with Ragin's work on 'multiple conjectural causation' (1987, 2000) as it implies that labour market institutions do not have a consistent causal effect on unemployment that applies universally. The fact that our sensitivity analysis and, in particular, changes in the sample lead to very different conclusions about the effects of institutions supports this interpretation. Overall, while our analysis does not find strong evidence about the adverse effects of labour market institutions on unemployment in general, we cannot exclude the possibility that institutions may be responsible for high unemployment in some countries. However, based on our analysis, there is little merit in recommendations that call for across-the-board institutional deregulation.

Conclusion

This paper has examined the role of labour market institutions in the determination of unemployment in the EU and OECD countries over the last three decades. The analysis entailed a re-examination of a number of specifications that have been used frequently in the empirical literature on unemployment which underpins the view that deregulation improves labour market performance. Our analysis pays special attention to common, but often neglected, problems associated with macro-comparative time-series cross-section analysis, such as the potential sensitivity of the results to the choice of estimators and small changes in model specifications and the sample. Robustness checks reveal that most results are fragile and that even a small number of observations may exert inordinate leverage on the coefficient estimates.

On the whole, we find no systematic support for the conventional view that unemployment is a consequence of rigid labour market institutions. Among the institutions only wage bargaining coordination shows a fairly robust association with unemployment, but this variable has a beneficial rather than a detrimental effect. Union density and benefit generosity, which showed up as significant in the initial analysis, do not survive the robustness checks. Turning to the interaction models, our initial analysis finds some indications that the complementary tax and benefit reforms may be beneficial, but these findings are less robust to changes in the sample. A re-estimation of the main models for the advanced and new market economies separately reveals that institutions play a more important role in the advanced economies, with several of them showing some association with unemployment. However, these initial results are also fragile to changes in specifications and the sample. In the advanced economies, union density is the only institutional variable that remains robustly associated with unemployment. The interaction between the benefit replacement ratio and the tax wedge also seems to play a role, though this association is more specification dependent. At the same time, in CEE no institution alone is directly and robustly associated with unemployment, but there are some (albeit not strong) indications that the interaction between the tax wedge on the one hand, and employment protection and the benefit replacement rate on the other may play a role. We also do not find sufficiently robust evidence for the hypothesis that institutions affect unemployment indirectly by amplifying the adverse effects of economic shocks.

In sum, our models provide no compelling evidence about the adverse effects of institutions. Meanwhile, GDP growth, and to a lesser extent the terms of trade, seem to be more consistent predictors of unemployment. Our analysis therefore challenges the policy orthodoxy that comprehensive labour market deregulation is necessary to reduce or stabilize unemployment. Calls for further deregulation seem to be especially unwarranted in CEE countries, where the link between institutions and unemployment appears particularly weak. But even in the advanced economies the effects of institutions are weak and in most cases depend heavily on which countries are included in the analysis. Given the lack of robustness, the most plausible interpretation of our results is that institutions have different effects in different contexts, and that therefore there is no universal cause of (and remedy for) unemployment. An implication of this for further research is that in-depth analyses of individual countries or particular groups of countries may be more fruitful in offering sound policy recommendations than the continued search for universal causes of unemployment. Of course, it is possible that the lack of a strong link between institutions and unemployment that we found may reflect the fact that institutions have different effects for different groups of the labour force. Clearly, our analysis of aggregate unemployment cannot identify such effects and further research on these issues is needed. Nonetheless, our findings suggest that any results from this type of quantitative macro-comparative research must be taken with a grain of salt and should not be used as unquestionable evidence for reforming particular institutions in a particular country.

Acknowledgements

We are particularly grateful to Lucio Baccaro, Barry Reilly, Bernhard Kittel, Lane Kenworthy and three anonymous reviewers for detailed and constructive comments. An earlier version of this paper was presented at the 23rd Annual Conference of the Society for the Advancement of Socio-Economics (SASE) and a fellow lecture series at the Hanse-Wissenschaftskolleg in June 2011.

Funding

Research for this paper was supported by the Economic and Social Research Council (ESRC) grant RES-061-25-0354.

Supplementary material

Supplementary material is available at *SOCECO* online.

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