research paper series

Globalisation, Productivity and Technology Programme

Research Paper 2022/08

Employment to output elasticities and reforms towards flexicurity: Evidence from OECD countries

Holger Görg, Cecília Hornok, Catia Montagna and George E Onwordi

Employment to Output Elasticities & Reforms towards Flexicurity: Evidence from OECD Countries*

Holger Görg† Cecília Hornok‡ Catia Montagna§ George E. Onwordi¶

September 16, 2022

Abstract

Labour market reforms in the direction of ’flexicurity’ have been widely endorsed as a means to increase an economy’s ability to adjust to negative shocks while offering adequate social safety nets. This paper empirically examines how such reforms influence employment’s responsiveness to output fluctuations (employment-output elasticity). To address this question, we employ a single equation error correction model with policy interactions on a panel of OECD countries, which also incorporates the period of the Great Recession, and distinguish between passive and active labour market policy types. Flexicurity is represented by three policy measures: unemployment benefit generosity, the flexibility of hiring and firing rules, and spending on active labour market policies. We find that the effects of any single policy change are shaped by the broader existing policy mix within which it takes place. A hypothetical flexicurity reform towards the policy mix of Denmark, a well-known example of the flexicurity regime, is found to increase or leave unchanged countries’ short-run employment-output elasticities, depending on the initial policy mix. These results are robust to accounting for a large set of additional labour market institutions.

Keywords: employment-output elasticity, labour market policy, welfare state, flexicurity
JEL codes: E24, E32, J21, J65

*This work was supported by the NORFACE ERA-NET (New Opportunities for Research Funding Agency Cooperation in Europe Network) Welfare State Futures Programme, Grant Number 462-14-120. We thank Nils Braakmann, Sebastian Könings, Steve Matusz, Fredrik Sjöholm, John Skåtun, Allan Sørensen, participants of the NORFACE Welfare State Futures Programme “Globalisation, Labour Markets and the Welfare State” final workshop and two anonymous referees for their valuable comments.

†Kiel Centre for Globalization, Kiel Institute for the World Economy and University of Kiel, Germany. E-mail: Holger.Goerg@ifw-kiel.de

‡Kiel Centre for Globalization and Kiel Institute for the World Economy, Germany. E-mail: Cecilia.Hornok@ifw-kiel.de

§University of Aberdeen and the Scottish Institute for Research in Economics, Scotland, UK. E-mail: c.montagna@abdn.ac.uk

¶University of Aberdeen, Scotland, UK. E-mail: george.onwordi@abdn.ac.uk

1 Introduction

The response of employment to external shocks is a central concern to policy-makers. Crivelli, Furceri and Toujas-Bernate (2012) show that unemployment rose substantially in advanced countries after 2008, while this did not happen in Latin American or African economies. In particular, they find that the employment responses to output shocks (output elasticity of employment) differ significantly among advanced countries – point estimates for the long-run elasticity range from 0.64 for Western Europe to 0.81 for North America. A number of papers looked at the role of institutional differences in determining these inter-country variations in the employment responsiveness to output shocks (Crivelli, Furceri and Toujas-Bernate, 2012; Bassanini, 2012; Blanchard and Wolfers, 2000).[1]

A high degree of inter-country heterogeneity in the types of institutions governing the labour market persists in the EU and indeed in the OECD. However, in recent years the flexicurity system – combining fairly generous unemployment support systems with flexible hiring and firing rules and high levels of active labour market policies to foster participation and employability – has been widely endorsed as a means to increasing an economy's ability to adjust to negative shocks while offering adequate social safety nets.[2] This support has in part been driven by the observation that, in the aftermath of the Great Recession, whilst many countries experienced massive increases in unemployment and output contractions, a flexicurity country such as Denmark suffered relatively minor unemployment consequences (Andersen, 2015). Surprisingly, however, empirical research that sheds light on the aggregate impact of flexicurity on labour market outcomes is limited, with extant studies focusing either on qualitative assessments (e.g., Andersen and Svarer, 2007) or on the effect of specific aspects of the flexicurity package (e.g., Card, Kluve and Weber, 2010; Faccini and Bondibene, 2012; Kluve, 2010). This is an important limitation since the relevance of the interaction between different policy instruments has been forcefully demonstrated theoretically – e.g. Davoine and Keuschnigg (2015), Dabusinskas, Kónya and Millard (2016) and Molana, Montagna and Onwordi (2021) – and, thus, matters for any country that aims to adopt the prescribed labour market reforms by institutions such as the IMF and the European Commission.

A first objective of this paper is to offer more up-to-date evidence, including the time of the Great Recession, on the employment-output elasticity and its changes – thus capturing the adjustments in employment and their level of persistence over time. Specifically, one of our paper's primary contributions is to take a step towards overcoming the relative lack of empirical evidence on the aggregate labour market effects of flexicurity reforms by focusing on their effects on the employment-output elasticity – which is an important indicator of how an economy responds to shocks. In addition, and most importantly, it will not only pay particular attention to the role of passive and active labour market policies (PLMP and ALMP, respectively) but also to their interactions in driving outcomes. The conjectures that guide our approach are based on the theoretical results of Molana, Montagna and Onwordi (2018) who demonstrate that different combinations of policy instruments that characterise

[1] At a theoretical level, the response of employment to output fluctuations may be driven by a number of factors including variations in the degree of capital mobility (Azariadis and Pissarides, 2007; Molana, Montagna and Onwordi, 2021), degree of openness to international trade (Cacciatore, 2014), the nature of the distribution of firm-size (Görg *et al.*, 2017), the structure of labour market institutions and active labour market policies characterising different economies (Molana, Montagna and Onwordi, 2018).

[2] Flexicurity policies are central to the European 2020 employment strategy and have, more broadly, been supported by international institutions such as the IMF (see, e.g., Blanchard, Jaumotte and Loungani, 2014).

flexicurity yield different degrees of volatility in employment/unemployment outcomes and highlight the importance of the initial conditions.

Our analysis covers both the Great Recession and the recovery period in a panel of 20 OECD countries. In particular, we consider two different time spans: from 1960 to 2012 and from 1986 to 2012; the latter enables us to explore the individual and combined effects of policy reforms. We first estimate the employment-to-output elasticity and find that, on average, the response of employment to output fluctuations in the short run increased substantially in the 1980s and 1990s, but has stalled since then. In contrast, the long-run elasticity has remained stable over time.

We then characterise labour market regimes by three policy measures that capture two typical PLMPs (i.e. the generosity of unemployment protection and the flexibility of the labour market with respect to hiring and firing) and an ALMP such as the expenditure on training programmes to enhance employability. When examined individually, we find that some specific policies have had substantial effects: both the generosity of unemployment benefit and higher degrees of flexibility in hiring and firing regulations induce greater short-run output elasticity of employment. The combined impact of these two policies is also found to have a sizeable offsetting effect on the responsiveness of employment to output fluctuations, implying the existence of complementarity of policy outcomes.

To further explore this complementarity effect, we consider three forms of labour market regimes – *average*, *liberal* and *interventionist*. The first is a hypothetical system that we obtain by setting the policy variables to their sample average values. The *liberal* regime, such as a liberal welfare state, is characterised by a high flexibility index (at the 75th percentile of the sample distribution) and relatively low levels of the other two policy variables (unemployment benefit and training expenditure), at their 25th percentiles. The *interventionist* regime, in contrast, is characterised by a relatively low flexibility index (at its 25th percentile) and relatively high values of the other two instruments (at their 75th percentiles). Our interest is to investigate the effects on the responsiveness of employment in the different regimes of implementing reforms in the direction of flexicurity. Consistent with the theoretical predictions of Molana, Montagna and Onwordi (2018, 2021), we find that policy complementarities matter when considering the impact of changes to individual policy instruments. For example, our estimations reveal that, ceteris paribus, raising the generosity of unemployment benefit in a *liberal* regime is not associated with higher short-run employment-output elasticity. In an *interventionist* regime, a reform towards more labour market flexibility is not accompanied by greater short-run employment responsiveness.

The above results suggest that reform 'packages' that shift the policy mix characterising each regime towards the Danish *flexicurity* model are not likely to have the same impact on the responsiveness of employment to output variations across all benchmark systems. In particular, further estimations reveal that for countries such as the US and Canada with liberal regimes, the flexicurity reform does not have statistically significant effects on employment's responsiveness. However, we find that labour market reforms toward flexicurity would imply greater short-run employment-output elasticities in, e.g., Australia, Japan, Germany and Great Britain in the order of approximately 0.2-0.25 percentage points. Interestingly, while Germany is both less interventionist and less liberal compared to Denmark (our benchmark) and has high levels of expenditure on training programmes, Great Britain is much less interventionist and more liberal, and has a lower level of active labour market policies. A key message of this paper, therefore, is that a flexicurity reform, as widely advocated, may result in very different aggregate effects depending on the characteristics of the initial regimes, especially in the short-run, resulting in greater employment volatility by amplifying the magnitude of the response to output fluctuations.

This paper contributes to the empirical literature that investigates the role of labour market institutions and policies on labour market outcomes (Bassanini, 2012; Bertola, Blau and Kahn, 2002; Blanchard and Wolfers, 2000; Crivelli, Furceri and Toujas-Bernate, 2012; Faccini and Bondibene, 2012; Nickell, Nunziata and Ochel, 2005). However, some of these works only look at the long-run impact of these policies on the labour market – with the exception of Bassanini (2012) and Faccini and Bondibene (2012) who consider the short-run dynamic implications. We look at both the short-run and the long-run elasticities but, unlike all of these papers, we also examine the complementarities of PLMPs and ALMPs from the perspective of a flexicurity reform using recent data evidence.

This study is also related to papers evaluating the macroeconomic effects of labour market policies. Gnocchi, Lagerborg and Pappa (2015) and Abbritti and Weber (2010; 2018) examine the effects of labour market institutions in driving business cycle fluctuations. Fonseca, Patureau and Sopraseuth (2010) also explore the relationship between labour market institutions and business cycles, focusing on international co-movements. The effects of institutions on the dynamic adjustment of inflation have been studied by Bowdler and Nunziata (2007), while Rumler and Scharler (2009) examine how institutions affect the volatility of output and inflation. We differ in our focus (and thus contribute to the debate) on the role of flexicurity reforms on aggregate labour market outcomes. From a broader context, a number of papers have assessed the benefits (or the effectiveness) of flexicurity reforms, but these studies are mostly qualitative: e.g., Andersen and Svarer (2007) and Andersen (2015) provide an assessment of the Danish experience, focusing respectively on workfare policies and long-term unemployment. Qualitative assessments of different welfare state reforms in selected countries in the EU have also been documented by Eichhorst and Konle-Seidl (2006) and Räisänen *et al.* (2012). By contrast, Kluve (2010) provides a quantitative assessment of the effectiveness of ALMP measures, concluding that training programmes have had mild effects on employment outcomes. Shahidi *et al.* (2016) assess the health-related impact of temporary contracts that may result from the less stringent hiring and firing regulations associated with flexicurity reforms.

In the next section, we present our empirical approach. Section 3 focuses on the description of the data and its time series properties. Section 4 obtains the baseline elasticity estimates and presents temporal changes and cross-country heterogeneity in the short-run elasticity estimate. Section 5 continues with a presentation of the selected labour market policies. Section 6 reports the estimates with the policy variables and discusses the effects of policy changes under different labour market regimes. This is followed by the evaluation of the flexicurity reform in Section 7. Finally, Section 8 presents some robustness checks, and Section 9 concludes.

2 Empirical approach

Taking advantage of the relatively long time dimension of our data, we follow standard time series methodology and describe the dynamic relationship between employment and output in a single equation error correction model (Hendry, 1995). This one-step estimation approach has also been popularised recently by Rao, Singh and Kumar (2010) and Kumar and Rao (2012) for panel data applications.[3] An advantage of the approach is that it encompasses the comovement of employment and output in both the short and long term and thus allows for the joint estimation of the short- and long-run employment to output elasticities. Related empirical literature typically estimate either short- or long-run elasticities only. Bassanini (2012), for instance, estimates elasticities for the short run by relying on cyclical fluctuations of the variables. In contrast, Kapsos (2006) or Crivelli, Furceri and Toujas-Bernate (2012) estimate long-run elasticities by regressing the level of employment on the level of output.

The error correction model between employment (e) and output (y), which we apply in this paper, is

$$\Delta e_{i,t} = \alpha \left[e_{i,t-1} + \beta y_{i,t-1} + \delta t\right] + \gamma \Delta y_{i,t} + \epsilon_{i,t}. \tag{1}$$

Equation (1) implies that the change in employment (annual and logarithmic) in country i and year t is driven by the change in output ($\Delta y_{i,t}$) in that country and year, the deviation of employment from its long-run equilibrium in $t-1$ (term in brackets) and a random error term ($\epsilon_{i,t}$) capturing contemporaneous employment shocks. The term in the square bracket is the long-run equilibrium relationship between employment and output, where β is assumed to be negative and it measures the long-run employment to output elasticity. The long-run relationship also allows for a deterministic linear time trend (δt), which may represent long-run technological or demographic changes that are common across countries. The short-run employment-output elasticity is captured by γ.

The model assumes that the time series of both employment and output are first-order integrated, so that their annual changes are stationary. It also anticipates that employment and output form a long-run cointegrating relationship of the form captured in the bracket. Empirically, this cointegrating relationship may or may not exist. If it exists, then $\alpha < 0$ in the model, which implies that any deviation in employment from its long-run equilibrium will result in an error-correction adjustment, where the speed of adjustment is captured by α.

This empirical approach, which does not rely on estimating aggregate labour demand or labour supply equations separately, enables us to investigate how employment fluctuates in response to output variations over time, regardless of whether this response originates from demand-side or supply-side shocks and of how the movements in output propagate through the labour market to affect employment. Our approach is deliberately atheoretical and can be perceived as a generalisation of Okun's Law – a purely empirical relationship. Of course, as with Okun's Law, the relationship between output and employment may also depend on other factors, such as labour productivity growth, changes in working hours per employee, or the size of the labour force. To account for these forces, we will present several robustness checks to our baseline empirical model.

[3] Gnocchi, Lagerborg and Pappa (2015) employ a similar approach to examine the direct relationship between labour market policy instruments and the business cycle. A similar empirical approach is taken for the estimation of the employment-output elasticity for the Gulf countries and for Turkey, respectively, by Behar (2017) and by Sahin, Tansel and Berument (2015) and for estimating the effects of short-time work schemes on various labour market outcomes by Hijzen and Martin (2013).

We reformulate equation (1) into the following regression equation,

$$\Delta e_{i,t} = \alpha e_{i,t-1} + \beta^* y_{i,t-1} + \gamma \Delta y_{i,t} + \delta_t + \delta_i + \varepsilon_{i,t}, \tag{2}$$

where α remains the speed of adjustment parameter, γ is the short-run elasticity, while the long-run elasticity is obtained as $-\frac{\beta^*}{\alpha}$. Estimating this equation instead of (1) allows for greater flexibility in the treatment of the common time trend, as it assumes year-specific fixed effects (δ_t) instead of a linear time trend. Moreover, the equation accounts for country-specific constants (δ_i). As discussed later in this paper, as part of robustness checks, we experiment with alternative specifications – for example, we add country-specific linear time trends to account for the possibility that secular trends, e.g. in terms of demography, productivity or technology, differ from country to country. As a more direct test for the role of demographic forces, in another specification, we estimate equation (2) with both employment and output divided by the size of the working-age population, a specification which directly accounts for changes in the size of the labour force. Furthermore, in yet another specification, we measure employment with full-time equivalent employment instead of the number of employees, thereby allowing for employment adjustments both at the extensive and intensive margins. The conclusions based on (1) remain robust regardless of the alternative specifications.

In the above framework, it is possible to assess the existence of a cointegrating relationship by testing whether α is significantly smaller than zero (Ericsson and MacKinnon, 2002). Should the estimated value of α be non-negative, then the cointegrating relationship would be considered non-existent, in which case equation (2) would simplify to

$$\Delta e_{i,t} = \gamma \Delta y_{i,t} + \delta_t + \delta_i + \varepsilon_{i,t}, \tag{3}$$

and the short- and long-run elasticities would both be γ. In what follows, we refer to equation (2) as the error correction (EC) specification and equation (3) as the first-difference (FD) specification.

To incorporate the role of policies we follow on previous literature (e.g. Blanchard and Wolfers, 2000; Crivelli, Furceri and Toujas-Bernate, 2012; Bassanini, 2012; Hijzen and Martin, 2013; Abbritti and Weber, 2018) and include interactions of the right-hand side variables with the policy variables in the regression equation. The FD specification will then become

$$\Delta e_{i,t} = \gamma_1 \Delta y_{i,t} + \gamma_2 \Delta y_{i,t} \tilde{P}_{i,t-1} + \eta \tilde{P}_{i,t-1} + \delta_t + \delta_i + \varepsilon_{i,t}, \tag{4}$$

where $\tilde{P}_{i,t-1}$ denotes the vector of policy variables (and possibly their interactions with each other), all lagged by one year. For ease of interpretation, the policy variables are 'centered' within each year, i.e. expressed as deviations from their country means as $\tilde{P}_{i,t} = P_{i,t} - \frac{1}{N}\sum_{j=1}^{N} P_{j,t}$. The short-run elasticity of any given country can thus be interpreted as an average elasticity plus country deviations originating from policy differences at any given point in time,

$$\text{sre}_i = \gamma_1 + \sum_k \gamma_2^k \tilde{P}_i^k \tag{5}$$

where k is the index for the different policy measures and – possibly – their interactions. The estimate for γ_1 will always give the short-run elasticity for the average country.[4] The estimated γ_2 will in turn measure how the policy variables (or their interactions) correlate with this elasticity or, to put it differently, what change in the elasticity associates with a one-unit increase in a given policy variable.

The EC specification (2) can similarly be extended to incorporate the policy variables. Specifically, doing so will result in estimating an equation such as (4) but augmented with the terms $\alpha_1 e_{i,t-1} + \alpha_2 e_{i,t-1}\tilde{P}_{i,t-1}$ and $\beta_1 y_{i,t-1} + \beta_2 y_{i,t-1}\tilde{P}_{i,t-1}$. This specification has the advantage of enabling us to also see how the policy variables influence the long-run elasticity or the speed of adjustment to the equilibrium. Similar to the short-run one, the long-run elasticity can be expressed as a function of parameters common to all countries and policy-induced individual deviations, more formally,

$$\mathrm{lre}_i = -\frac{\beta_1 + \sum_k \beta_2^k \tilde{P}_i^k}{\alpha_1 + \sum_k \alpha_2^k \tilde{P}_i^k}. \tag{6}$$

Policy variables can, of course, relate endogenously to simultaneous employment dynamics. It is natural to assume that, in some cases, national governments adjust their labour market policies to employment shocks. In order to alleviate such reverse causality issues, we use the first lag of the policy variables in our regressions.[5] Moreover, in order to lessen endogeneity concerns from omitted variables, we control for several country-time-specific confounders in the regressions. Possible confounders are, e.g., other labour market indicators, such as union density or the tax wedge, but also the degree of openness to international trade or the size of the government. These can explain cross-country differences in the employment-output relationship and, at the same time, correlate with the labour market policy mix. Of course, to the extent that our observable variables, and our one-period lag, do not capture all possible endogeneity issues, our results may be somewhat biased. This needs to be kept in mind in the interpretation of results. Some recent empirical studies on the effect of certain labour market policies on employment use more refined identification strategies in that they identify from major reform events or rely on industry-level variation in employment responsiveness (e.g., Duval, Furceri and Jalles, 2020; Ciminelli, Duval and Furceri, 2020) or use instrumental variable estimation techniques (e.g., Hijzen and Martin, 2013). While these strategies are undoubtedly better able to infer causal relationships, the studies mentioned above focus only on individual policies. Our interest in the labour market policy *mix* as well as our focus on both the short- and long-run horizons limit the possibilities for such estimation strategies.

3 Data and time series properties

Our database is an annual panel of 20 OECD countries. We start with a balanced panel of employment and output for 53 years (1960-2012). As a result of the subsequent inclusion of the policy variables in the analysis, the time span reduces to 27 years (1986-2012) and the panel becomes slightly unbalanced. The set of countries consists of thirteen pre-2004 European Union members (not including Luxembourg and Greece[6]), Australia, Canada, Japan, New Zealand, Norway, Switzerland, and the US. We source

[4] This is ensured by centering the policy variables.

[5] Abbritti and Weber (2018) also use the lags of the policy variables in their interacted panel VAR model to partially correct for possible endogeneity issues.

[6] Luxembourg and Greece are exluded because important labour market indicators are either not available for them or only available for a limited time span.

the time series for employment and output from the Penn World Table (PWT, version 10.0).[7] Output is measured by the real GDP at constant 2017 prices (in million US dollars), while employment is captured either by the number of persons employed or, in an alternative specification, the full-time equivalent (FTE) employment. We calculate FTE employment as the number of employees multiplied by the average annual hours worked per employed person and divided by the total number of hours a full-time employee works in a year (40 hours $\times$ 52 weeks). We relegate the description of the labour market policy variables to Section 5, while further definitions and data sources are summarised in Table A.1 in the Appendix.

Before we delve into the econometric analysis, we consider the time series properties of employment and output. In particular, we carry out unit root tests on the panel of 53 years both country-by-country (ADF, KPSS) and panelwise (Levin-Lin-Chu, Fisher-ADF and Hadri). The results of these tests are available in Tables A.2 and A.3 for the country-by-country tests and in Tables A.4 and A.5 for the panel. Because the results are qualitatively similar for the two employment measures, we only report the results for the number of employees series. Based on these test results we conclude that, in levels, both employment and output display unit roots against being trend stationary, while their first differences are stationary for the majority of the countries. A notable exception is Spain, where the results cast doubt on whether the first difference in employment can be considered stationary.

More specifically, the country ADF tests cannot reject the null hypothesis of a unit root for the levels (except for output in Switzerland and Spain), while they always reject it for the first differences (except for employment in Spain). The KPSS test rejects the null of stationarity for the levels in most of the cases, while the results for the first differences are somewhat mixed, as they reject stationarity in quite a few cases. Because the KPSS test is known for rejecting too often, i.e. having a high rate of type I error, we rely here more on the ADF test. The panel unit root tests largely confirm that both series have a unit root in levels. As for the first differences, the Levin-Lin-Chu and Fisher-ADF tests clearly reject the null of a unit root. The Hadri test rejects the null hypothesis, indicating that, for some countries, the first differences cannot be considered stationary against alternative hypothesis of some panels having unit roots. These results are similar if we run the tests on a panel excluding Spain.

With these results in mind, we continue with the assumption that output and employment have unit roots, while their first differences are stationary. However, to account for the mixed results of the unit root tests for Spain, most results in this paper are reported both with and without Spain in the sample.

Next, to assess the existence of cointegrating relationships between employment and output, we apply two types of tests. First, we perform the Johansen cointegrating rank test on each country (Johansen, 1991), while allowing for a time trend in the cointegrating equation. The resulting trace and maximum-eigenvalue statistics are reported in Table A.6, together with the relevant 5% critical values at the bottom of the table. For the majority of the countries, we find a cointegrating equation between employment and output. When the number of cointegrating equations is chosen by minimising the HQIC (Hannan and Quinn Information Criterion), we find cointegration for all the 20 countries. Second, in the next section, we test for the presence of cointegration by examininig the statistical significance of the speed of adjustment parameter when estimating equation (2) on the pooled

[7]The Penn World Table we use is described in Feenstra, Inklaar and Timmer (2015). The data is available for download at `www.ggdc.net/pwt`.

data. A negative and significant estimate would confirm the existence of cointegration. Because the conventional critical values for statistical significance do not apply under the null hypothesis of no cointegration, we use the critical values produced by Ericsson and MacKinnon (2002) and refined by MacKinnon (2010). As the results in Table 1 in the next section and Table A.7 show, the t-statistics for the speed of adjustment parameter estimate (in squared brackets) lie outside the 5% critical values, indicating the existence of cointegration.[8]

4 Elasticity estimates

How responsive was employment to output fluctuations in the past decades in developed OECD countries? To what extent does this relationship vary over time or across individual countries? Assuming a permanent output change, does an immediate response differ substantially from the long-run effect? To answer these questions, this section presents estimates of the short-run and long-run employment to output elasticities and explores time and country variations therein.

4.1 Baseline estimates

We estimate the average value of the employment to output elasticity for our panel of OECD countries. Using the empirical model (2) we can identify both the short-run and the long-run estimates in one step and test for the existence of a cointegrating relationship. For robustness and for reasons discussed above, we also report the elasticity estimates from model (3). We consider both the full time span (1960-2012) as well as a shorter panel (1986-2012) which corresponds to the time period for which labour market policy variables are observed. Further, we report results both with the number of employees and the FTE employment as employment variables. All regressions include a full set of country and year-fixed effects and are estimated with Ordinary Least Squares and robust standard errors. Table 1 reports the results.

The short-run employment-output elasticity of the average country in our sample is estimated to be 0.3-0.4 on the long panel and 0.5-0.6 on the shorter one. An estimate of 0.5, for example, indicates that a 1% increase in the volume of output is accompanied by a 0.5% instantaneous increase in employment. The fact that our estimate on the shorter – and more recent – panel is larger than the estimate on the full panel shows that, for the average country, employment has become more responsive to output shocks over the recent decades. The estimate of the long-run employment-output elasticity, obtained from the regression with the number of employees and using the EC specification, is around 0.8 for the average country. This number implies that if output permanently increases by 1%, the number of employees will rise by 0.8% in the long run, ceteris paribus.[9] We note, however, that the long-

[8]Note that both methods we use to test for cointegration are developed for single country time series. Although there exist cointegration tests for panel data (Pedroni, 1999; Westerlund, 2007), these have been developed for large-T (and at least medium-N) panels and their reliability on panels of limited size in both dimensions is not yet proven (Behar and Hodge, 2008). Based on the Pedroni and the Westerlund panel tests we fail to reject the null of no cointegration. With this consideration in mind, we proceed reporting estimates for both the EC and the FD models and indicate whenever the presumption of cointegration leads to qualitatively different results.

[9]Figure A.1 in the Appendix, which was produced by making dynamic forecasts of our estimated model with and without a 1% shock to output and taking the difference of the two forecast paths, illustrates this dynamic adjustment of employment to a 1% permanent output shock over time. As is evident from the figure, the instantaneous impact is greater in the shorter panel than in the longer one. However, the adjustment of employment as the impact of the

Table 1: Elasticity estimates

Depvar: Δe	Number of employees				FTE employment			
	1960-2012		1986-2012		1960-2012		1986-2012	
Δy (SR elasticity)	0.317***	0.343***	0.521***	0.515***	0.408***	0.428***	0.590***	0.594***
	(0.028)	(0.027)	(0.045)	(0.041)	(0.031)	(0.031)	(0.054)	(0.050)
e_{-1} (Speed of adj)		-0.034***		-0.084***		-0.035***		-0.083***
		(0.004)		(0.015)		(0.004)		(0.013)
		[-8.26]		[-5.71]		[-8.63]		[-6.32]
y_{-1}		0.027***		0.061***		0.020***		0.046***
		(0.004)		(0.012)		(0.006)		(0.015)
Observations	1,040	1,040	540	540	1,028	1,028	540	540
R-squared	0.477	0.517	0.621	0.655	0.495	0.525	0.627	0.660
LR elasticity		0.781***		0.726***		0.584***		0.550***
		(0.134)		(0.101)		(0.151)		(0.136)
CI test 5% crit val		-3.97		-4.15		-3.97		-4.15

Notes: Estimation of models (2) and (3) are based on OLS, where employment is either the number of employees or fulltime equivalent (FTE) employment. Country and year dummies are included in all regressions. Parentheses include robust standard errors, brackets t-statistics. Standard errors for the long-run elasticity are obtained with the delta method. The CI test is the Ericsson-MacKinnon cointegration test, with null hypothesis of no cointegration. Critical values (obtained from MacKinnon, 2010) must be compared with the t-statistic for the speed of adjustment parameter estimate. We can reject the null of no cointegration. *** p<0.01, ** p<0.05, * p<0.1.

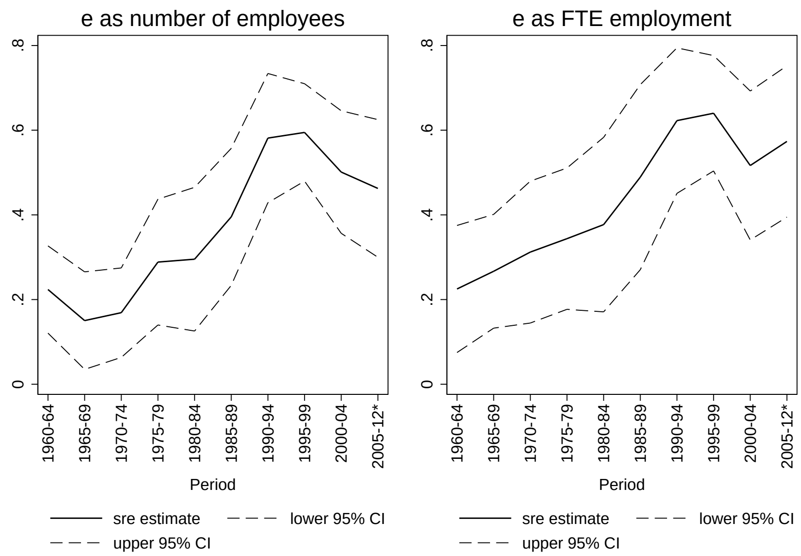
run elasticity is estimated with large uncertainty, as indicated by its 95% confidence interval ranging between 0.5 and 1. When FTE employment is used, we obtain point estimates that are somewhat higher for the short-run elasticity and lower for the long-run elasticity. Notice, though, that the confidence intervals around these estimates overlap, meaning that the two types of estimates are not different statistically. Nevertheless, the direction of the differences is economically sensible. In the short run, some of the employment adjustment is likely to take place on the intensive margin because adjustment on the extensive margin (hiring/firing) is more costly. If there is a negative output shock, employees may temporarily work reduced hours. Similarly, if the output shock is positive, existing employees may initially be required to work overtime before new ones are hired. In the longer run, however, if the shock turns out to be permanent, extensive margin adjustment will also take place. In fact, the long run adjustment on the extensive margin may be larger than for FTE employment if part-time employment becomes more widespread over time, which has been the case in the labour markets of most OECD countries in recent decades.

Our elasticity estimates are roughly consistent with those found for developed countries in previous literature (e.g. Kapsos, 2006; Crivelli, Furceri and Toujas-Bernate, 2012), despite the different estimation methodologies and time series length. These studies also look at a larger set of countries, including less developed economies, and show that developed countries typically have larger employment-output elasticities than less developed ones.

Results reported thus far are based on the sample including Spain. Results excluding Spain are reported in Table A.7 in the Appendix; they are largely comparable to the estimates reported here.

increase in output dies out is the same for both panels, essentially resulting in similar long-run elasticities.

Figure 1: Short-run elasticity estimates over time



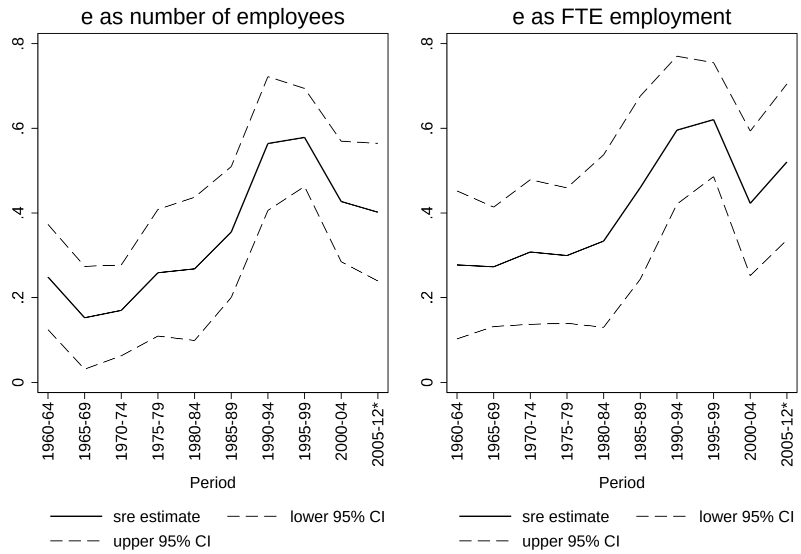
4.2 Estimates by period and country

Next, we investigate more closely how the short-run elasticity changes over time. To this end, we interact the output change variable on the right-hand side of equation (2) with dummies for five-year periods (except for the last period of 8 years) and estimate it on the long panel. The resulting period-specific elasticity estimates for both employment measures are plotted in Figure 1 for the sample of 20 countries and in Figure 2 for a sample without Spain.[10] The time pattern of the short-run elasticity is very similar across the different charts. It shows a clear upward trend until around the end-1990s, rising rapidly from around 0.2 (0.3 if FTE employment is considered) in the first two decades of the sample to around 0.6. After the end of the 1990s, however, the increase seems to have stalled or even reversed slightly.

Although there may be several factors underlying the observed time pattern of the short-run elasticity, labour market institutions and policies are certainly among the important ones. For example, as suggested by Blanchard, Bean and Münchau (2006), the sharp increase at around the middle of the sample period may be related to the introduction of more generous unemployment insurance systems in many OECD countries, as a response to large increases in the unemployment rate during the early 1980s. Similarly, the spread of the dual labour market across developed countries could have contributed to the upward trend in elasticity, as temporary employment contracts subject to less protective regulations enable stronger employment responses to output shocks (e.g., Bentolila *et al*, 2010). Furthermore, a labour market policy that could potentially contribute to the recent stabilisation/decline of the short-run elasticity are short-time work schemes, which have become a popular policy tool in OECD countries during the Great Recession (Hijzen and Martin, 2013). Short-time work programmes intend to preserve jobs by encouraging work sharing and by providing income support to

[10] Estimating the FD specification (3) results in very similar graphs.

Figure 2: Short-run elasticity estimates over time (without Spain)



workers whose hours are reduced, enabling a more muted response of employment to negative shocks.

Beyond a general trend in labour market institutions, the nature of the output shock and the type of countries affected by these shocks may also influence the strength of employment response. In particular, employment may respond to a positive output shock differently from a negative one because the rigidities in the labour market usually have an asymmetric nature (e.g. downward nominal wage rigidities). And since the degree of labour market rigidities varies from country to country, it also matters which country is affected by a shock. The literature on business cycle asymmetry documents that asymmetric labour market rigidities can amplify the business cycle in downturns, meaning that employment and output rise slowly in the expansionary and fall fast in the contractionary phases of the cycle (McKay and Reis, 2008; Abbritti and Fahr, 2013). It is also shown that in most European countries the degree of this asymmetry is larger for employment than for output, implying that the employment to output elasticity of these countries is larger during economic downturns than during expansions (Abbritti and Fahr, 2013; Kolasa, Rubaszek and Walerych, 2021).[11]

We use our empirical model to carry out a simple test for the presence of such asymmetries. First, we identify periods of economic slowdown for each country in our sample, where we rely on the literature on turning point analysis but with the important distinction that we have annual data.[12]

[11]The asymmetry found in Europe is consistent with Kolasa, Rubaszek and Walerych (2021), who find that - against the conventional view - hours per worker in European countries are much less procyclical than in the US, and in some economies even co-move negatively with output. This means that hours per worker may increase in Europe in times of negative shocks, aggravating the upward pressure on unemployment.

[12]We generate the GDP growth cycle for each country with the band-pass filter of Baxter and King (1999) and identify slowdown periods as years on the contractionary phases of the cycle. Contractionary phases are the years from each peak until the next trough. Harding and Pagan (2002) developed an algorithm to find peaks and troughs for quarterly data. We apply their method – using the user-written Stata command sbbq – but with altered parameters because of the annual frequency. We set the window over which local minima and maxima are computed as well as the minimum phase length to 1 year (instead of 2 quarters optimal for quarterly data), while the minimum cycle length is set to 2 years (instead of 5 quarters). We are aware of the problem that higher frequency information is lost when annual data

Second, we create a binary variable for the slowdown observations and include this variable and its interaction with output growth in the baseline regression. We run regressions for the whole sample as well as for a restricted sample of European countries, in order to check if the degree of asymmetry is indeed greater for Europe. The regression results reported in Table 2 suggest the existence of some degree of asymmetry. The point estimates for the interaction term are positive in all specifications – but are only statistically significant for the European subsample. Assessing the relative frequency of slowdowns in our sample by decade in Table A.8 we found that contractionary years were most common during the 1980s and 1990s. These findings indicate that some of the documented rise in the estimated short-run elasticity over these decades may be due to more frequent economic slowdowns causing larger employment responses in many of our sample countries.

Table 2: Business cycle asymmetry of the elasticity

Depvar: Δe	Number of employees				FTE employment			
	All countries		Europe only		All countries		Europe only	
Δy	0.331***	0.366***	0.355***	0.357***	0.392***	0.422***	0.369***	0.369***
	(0.040)	(0.039)	(0.045)	(0.043)	(0.046)	(0.046)	(0.051)	(0.051)
Slowdown	0.002	0.002	0.001	0.000	-0.001	0.000	-0.002	-0.002
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
$\Delta y\times$ Slowdown	0.049	0.049	0.118**	0.109**	0.088	0.087	0.152**	0.143**
	(0.053)	(0.049)	(0.057)	(0.054)	(0.056)	(0.053)	(0.063)	(0.061)
e_{-1}		-0.035***		-0.032***		-0.036***		-0.034***
		(0.004)		(0.008)		(0.004)		(0.008)
y_{-1}		0.027***		0.027***		0.021***		0.020***
		(0.004)		(0.005)		(0.006)		(0.007)
Observations	1,040	1,040	780	780	1,028	1,028	778	778
R-squared	0.483	0.524	0.517	0.542	0.499	0.530	0.504	0.520

Notes: Estimation of models (2) and (3) are based on OLS, where employment is either the number of employees or fulltime equivalent (FTE) employment. Slowdown is a binary variable which is 1 for country-years during the contractionary phase of the growth cycle. The cycle is computed using band-pass filter and peaks and troughs are selected based on Harding and Pagan (2002) but using annual series. Country and year dummies are included in all regressions. Parentheses include robust standard errors. *** p<0.01, ** p<0.05, * p<0.1.

Finally, we explore the extent of inter-country heterogeneity in the short-run employment to output elasticity. Specifically, we estimate model (2) by interacting the output change variable with country dummies to obtain country-specific elasticities. Figure 3 reports the short-run elasticities by country with 95% confidence intervals, estimated either on the long panel or the short panel, and plotted in descending order of the elasticity obtained from the long panel. Clearly, a considerable inter-country variation exists in the responsiveness of employment to output. Countries with the largest elasticities on the left-hand side of the graph (Spain, US, Ireland) have at least three times higher employment responsiveness than countries on the right-hand side (e.g., Italy, Austria). The point estimates typically get larger as the sample is shortened to the more recent decades; for most countries, however, these changes are not statistically significant, as the overlapping confidence intervals indicate. Exceptions are Spain and Portugal, whose short-run elasticities are significantly larger when estimated on the short (more recent) panel than on the long panel.[13] Spain is in a special position among the countries

is used. As it turns out, the solution provided by the Harding-Pagan algorithm under the above parameter values is trivial. The years of slowdown that we identify are exactly the same as if we simply take the years in which the GDP cycle declines relative to the previous year. Figure A.2 indicates periods of economic slowdown by country with grey color.

[13]These findings are largely unchanged (though with some changes in the country ranking) when estimating the

in that its estimated employment to output elasticity is among the highest irrespective of the length of the sample or how we measure employment. Moreover, the responsiveness of the number of employees to output shocks, as estimated on the more recent panel, is significantly higher than the elasticity of any other country in our sample. Bentolila *et al.* (2010) investigate the specificities of the Spanish labour market (in comparison to France) and explain the high responsiveness of employment to output fluctuations with the exceptionally high share of temporary employment contracts in Spain and the large existing gap between the dismissal costs of permanent versus temporary workers.[14]

Figure 3: Estimated short-run elasticities by country

5 Labour market policies in OECD countries

The remaining part of this paper investigates how labour market institutions and policies influence the employment-output elasticity. In particular, we focus on three policies which we identify as key elements of the flexicurity regime: (i) the generosity of unemployment benefits, (ii) the flexibility of the labour market with respect to hiring and firing, and (iii) the expenditure on training programmes to enhance employability. The first two are considered passive labour market policies, while the latter is a prime example of active labour market measures. Our interest in this set of labour market policy instruments is based on the fact that they reflect the key pillars of *flexicurity* – which has been widely endorsed as a means to reconciling the need for the flexibility required to adjust to negative shocks with that for adequate social safety nets (see, e.g., Blanchard, Jaumotte and Loungani, 2014). The descriptive statistics of the three policy variables – together with additional labour market indicators

first-difference specification.

[14]According to OECD data, no country included in our sample has ever had a higher share of temporary employment than Spain. For the average of the period 1990-2019, for which data on temporary employment is available, this share is 30% for Spain and only 10% for the rest of the countries.

that we discuss later – are reported in Table 3 below.

Table 3: Policy variables and their descriptive statistics

Variable	Full name	N	Mean	Std. Dev.	Min	Max
The three pillars of flexicurity						
grr	Gross unemployment benefit replacement rate	529	0.311	0.120	0.025	0.652
flex	Labour market flexibility index	529	4.055	0.965	1.677	5.746
train	Training expenditures per unemployed (log)	529	7.060	1.250	3.150	9.252
Additional policy variables						
uden	Trade union density	529	0.366	0.207	0.078	0.874
cove	Adjusted bargaining/union coverage rate	529	0.678	0.269	0.130	1.000
cent	Centralisation of collective bargaining index	529	2.480	1.047	0.875	5.000
taxw	Tax wedge	529	0.516	0.125	0.298	0.856
part	Prevalence of part-time employment	529	0.181	0.077	0.000	0.333
temp	Temporary employment share	365	0.121	0.064	0.037	0.350

Notes: Own calculations based on data sources listed in Table A.1. The tax wedge is from Abbritti and Weber (2018). The descriptive statistics refer to the period 1985-2011 and 20 countries including Spain. Training expenditures are adjusted for purchasing power parity differences and expressed in the natural logarithm of PPP US dollars.

The generosity of the unemployment benefit system is proxied by the gross replacement rate (*grr*), which is the ratio of the gross unemployment benefit level to the previous gross earnings of the average production worker in a country.[15] *grr* is a summary measure, calculated by the OECD as an average for two earnings levels, three family situations and three durations of unemployment. The series are available for every odd year between 1961 and 2011. To complete the time series, we fill in the even years with linear interpolation.[16] The benefit replacement rate of the average country in our sample is approximately 0.3, implying that the unemployment benefit is close to one-third of the previous gross earnings. The degree of benefit generosity varies considerably across countries: e.g., Denmark, the Netherlands and Norway have higher-than-average replacement rates, while Canada, Japan, Great Britain and the US provide weaker unemployment protection (Figure A.3). The variation over time is also non-negligible, with Italy having increased its replacement rate considerably in the first half of the sample and with Denmark and the Netherlands decreasing it somewhat in the more recent years.

Finding an explicit measure of labour market flexibility for both hiring and firing is empirically challenging. However – as is well demonstrated in the theoretical literature (see, for instance, Zanetti, 2011; Cacciatore, Fiori and Ghironi, 2016) – employment protection affects both job creation and job destruction. Hence, greater restrictions that are aimed at protecting workers from job losses, also lead to heightened disincentives to hire these workers. Thus, to capture flexibility in this paper, we adopt the inverse of the OECD Employment Protection Index (EPL), available for the years 1985-2013. The EPL is a composite index, calculated as a weighted average of the employment protection sub-indices

[15] We are aware that the net replacement rate, the ratio of net benefits to net earnings, would be a better measure for our purposes. The net rate is, however, produced by the OECD only since 2001 and hence would dramatically shorten the time series dimension of our study.

[16] The indicator was originally constructed for the OECD Jobs Study (OECD, 1994) and calculated for the average production worker. This series is available until 2005. Numbers for the more recent years are calculated for the average worker (both production and non-production) using the OECD Tax-Benefit Models. To minimize the impact of this methodological break, we rely on the former series until 2005 and extend it until 2011 by using the corresponding time changes in the latter series. This simple extrapolation naturally assumes that the dynamics of the two series are similar. Comparing the time changes of the two series for the years where both are available (2001-2005) we find that they indeed correlate strongly, with correlation coefficient around 0.9.

for the dismissals from regular contracts and the use of temporary contracts.[17] Our *flex* variable is the inverse of the EPL index ($flex = 6 - EPL$), ranging from 0 to 6, and showing how easy it is for employers to lay off regular workers (firing) or to employ them on temporary work contracts (hiring), as laid down by a country's regulatory system. Hence, a higher value of *flex* means a more flexible labour market, which entails both lower costs of hiring and firing. According to *flex*, the least flexible labour markets can be found in Southern Europe, while the most flexible ones are in the US, Canada, Great Britain and Ireland. Time series graphs in Figure A.4 also reveal a tendency among countries with less flexible labour markets to increase flexibility over time, which especially applies to Germany, Sweden, Spain, Italy and Portugal. Interestingly, Spain belongs to the least flexible countries in our sample based on *flex*, which is clearly at odds with the observed large flows of temporary workers into and out of employment in Spain (Bentolia et al., 2010). The reason is that the EPL index does not reflect the large gap between the dismissal costs between permanent and temporary workers. Therefore, we conclude that, despite its attractiveness and widespread use in the literature, the EPL index - and hence *flex* - is an inadequate indicator for the Spanish labour market.

We capture the importance of active labour market policies with the ALMP expenditures on training programmes per unemployed person (*train*). This category of training is defined to only include programmes (both institutional and workplace-based) that are *targeted* at unemployed persons or employees at high risk of becoming unemployed. Training programmes which are part of general education are excluded from this definition. Although the full spectrum of ALMPs includes a variety of different measures such as job creation, rehabilitation, or job sharing, we concentrate on training for two main reasons. First, we believe it best represents the motive behind the flexicurity idea of actively promoting workers' employability. Second, training makes up for the largest share of ALMP expenditures, totalling up to 45% for the average country-year in our sample. To make our *train* variable comparable across countries, we convert it to a common currency (US dollar) and adjust for purchasing power parity differences. Finally, we express it in natural logarithm. All data for the calculation of this variable are sourced from the OECD, with 1985 as the starting year. The sample average of *train* is 7.06, i.e. 1,164 PPP US dollars per unemployed person annually. As for cross-country differences, training expenditures turn out to be higher (and increasing) in European countries, as compared with non-European OECD members, which suggests the relatively high importance of ALMPs within Europe (Figure A.5).

The above three labour market policy measures (*grr*, *flex* and *train*) are obviously not applied in isolation but form parts of a policy regime. Cross-country policy correlations in Table A.9 reveal that countries with generous unemployment benefits tend to spend more on training schemes. This positive relationship is highly statistically significant and has become stronger over time. One possible reason for this is that generous unemployment benefit schemes need to be complemented with active labour market policies in order to sufficiently motivate the unemployed to seek work (Blanchard, Jaumotte and Loungani, 2014; Nickell, Nunziata and Ochel, 2005). Furthermore, the data also reveal that the generous systems tend to coexist with less flexible labour markets, though this relationship becomes statistically significant only in the more recent years of our sample period. In fact, along these three

[17] Following the OECD practice, $\text{EPL} = \frac{7}{12}\text{EPRC} + \frac{5}{12}\text{EPT}$, where EPRC is the sub-index for dismissals from regular contracts and EPT is the sub-index for the use of temporary contracts. A detailed description is provided in OECD (2013).

labour market measures, most countries in our sample can be broadly categorised into two regimes, one with a flexible labour market and low generosity (which we term 'liberal') and one with generous programmes and less flexibility ('interventionist'). The *liberal* regime mostly exists in Anglo-Saxon countries (especially in Canada, Great Britain and the US), the *interventionist* in most countries of continental Europe and Scandinavia.

Besides the three labour market policies that characterise the flexicurity aspect, our analysis also accounts for the role of further labour market institutions in influencing the employment to output elasticity. We aim at a relatively comprehensive representation of labour market institutions while acknowledging limitations due to data availability. Motivated by recent literature on the business cycle implications of labour market institutions (Abbritti and Weber, 2010; 2018; Gniocci *et al.*, 2015), we consider the following additional policy variables: the trade union density; the collective wage agreement coverage rate among employees; the degree of centralisation of the wage bargaining process (a summary index ranging between 0 and 5); and the tax wedge which measures the difference between the employer's labour costs and the net salary received by the employee.[18] Furthermore, we also account for the varying importance of non-fulltime and non-permanent employment across countries and over time. Because official OECD data on the share of part-time and temporary contracts is of limited availability, we experiment with an outcome-based measure. Using data on the average annual hours worked per employee (avh) from the Penn World Table and assuming that full-time annual working hours equal 2,080 (52 weeks $\times$ 40 hours per week) we generate a variable for the prevalence of part-time employment as $1 - \text{avh}/2080$. Note that this measure does capture not only the importance of part-time employment but also some temporary employment, namely employment shorter than a year (seasonal work). Finally, in a robustness check, we also consider OECD data on the share of temporary employment contracts, which however considerably reduces our sample size. Table 3 presents summary statistics of all these policy variables, while Table A.1 lists the data sources.

6 Elasticities and the labour market policy mix

In what follows, we investigate how individual policies and the policy mix relate to the responsiveness of employment to output shocks. Due to the aforementioned peculiarities of the Spanish labour market, we have decided to exclude Spain from the baseline estimation, but conduct a separate discussion and robustness check on it.

As our baseline, we estimate the EC specification while including the three flexicurity policy variables and their interactions with output growth and the lagged values of employment and output (Δy_{it}, y_{it-1}, e_{it-1}). For robustness, we also estimate the FD specification as in equation (4). The estimation results are reported in Table A.11 for the EC and Table A.12 for the FD specification. All the presented regressions include country-fixed effects, common year effects and a small set of country-time-specific control variables that can potentially affect employment elasticities. The latter includes the size of the working-age population, openness to international trade, and the size of the government (see a more detailed description in Table A.1).[19] Though not reported, these control variables turn out

[18]We are grateful to Mirco Abbritti and Sebastian Weber for generously providing us with their tax wedge variable used in Abbritti and Weber (2018).

[19]Panel unit root tests indicate that trade openness and government size are first-order integrated while the working-

to be significant predictors of employment growth in most regressions – while population growth and the size of the government associates positively with employment, the coefficient for trade openness is negative, though not quantitatively large. Furthermore, in the last two regressions we also account for the role of additional labour market indicators and their interactions with output growth and the lagged values of employment and output. We chose to include the first five of the six additional LMIs listed in Table 3, which is the largest possible selection without compromising the sample size. Similar to the three flexicurity policy variables, the LMI variables are centered and lagged by one year.

The estimates in Table A.11 can be interpreted as follows. The first row shows the short-run employment-output elasticity estimate for the average country, which is around 0.5 and robust across model specifications. The coefficients for the interactions of output growth with the policies capture how this elasticity changes if a country deviates in a policy measure from the average. The estimate of 0.843 (significant at 1% level) in the second row of column 4, e.g., means that, all things being equal, a country with an unemployment replacement rate that is higher than the average by 0.1 (10 percentage points) tends to have a 0.0843 higher-than-average short-run elasticity. Furthermore, a one-point higher flexibility index corresponds to a 0.113 higher elasticity (estimate of 0.113*** in column 4). These results show that greater unemployment benefit and flexibility in the labour market tend to amplify the short run effects of output changes on employment and, as we will discuss below, they are consistent with theoretical predictions. In contrast, training expenditure is not found to have a significant explanatory power for the short-run elasticity; this is likely due to the fact that training takes longer to have effects on workers' productivity. Note that the coefficient estimates in the different columns of the table are reasonably stable. Nevertheless, the inclusion of the additional LMIs and their interactions (20 additional regressors) in the last two columns is accompanied by a substantial loss of degrees of freedom in an already demanding empirical specification and causes some of the key estimates to lose their statistical significance.

Columns 5 and 7 include the results for all possible interactions between the flexicurity policy variables (*grr* x *train*, *grr* x *flex*, *train* x *flex*, and the triple interaction *grr* x *train* x *flex*). Broadly, the results show that policies do not only matter individually but also in combination. In particular, we find that although when taken individually, higher benefit generosity and more labour market flexibility are both associated with a higher employment responsiveness, there is an offsetting effect when the two are considered jointly. Testing the joint significance of the terms with policy interactions (test results are reported at the bottom of the table) also confirms that accounting for policy complementarities significantly improves the model fit. We conclude that the existence of such complementary effects among the policy variables indicates that the consequences of any single policy change is affected by the broader policy mix within which it takes place. Finally, the above results are largely unchanged when we estimate the FD specification (Table A.12).

The results above are likely to reflect the complex interactions and the feedback effects of policies on output-employment dynamics. Our findings are consistent with existing theoretical predictions. From the perspective of a search and matching framework, a higher generosity of unemployment benefit increases the responsiveness of employment to shocks (Molana, Montagna and Onwordi, 2018; Shimer,

age population is second-order integrated (Table A.10). This motivates us to include in the regressions the annual changes of trade openness and government size and the change in the annual change (second-order difference) of the working-age population.

2005; Zanetti, 2011) as our estimates reveal. The driving force is that a higher benefit raises the workers' outside option and their threat point in wage bargaining, which reduces match profitability. This in turn makes firms more 'sensitive' and implies that a shock to productivity will result in greater movement (in percentage terms) in terms of vacancy creation and employment. This result has been shown to hold whether job separation is exogenous (e.g., Hagedorn and Manovskii, 2008) or endogenous (Zanetti, 2011). The effects of greater labour market flexibility are, at a theoretical level, more nuanced. On the one hand, greater flexibility in hiring (e.g. lower vacancy creation costs) has been shown to induce higher employment responsiveness. On the other hand, increased flexibility in terms of firing rules can have a dampening effect on employment fluctuations, since it raises job match surplus, making firms less sensitive to a productivity shock. Thus, when both policies are implemented jointly (i.e. via increases in both hiring and firing flexibility), the impact on employment fluctuations will depend on which effect dominates. The positive impact on the short-run elasticity obtained in this paper suggests that empirically the flexibility in hiring regulation may have been the dominant force behind employment fluctuations. This is consistent with the fact that the strictness of firing regulation has been found empirically to have a mild (and even ambiguous) effect on firms job creation activities and thus employment (e.g. OECD, 2013). The offsetting effects of unemployment generosity and flexibility also produce an interesting result, which can also be explained from a theoretical perspective. First, as already established, the effects of higher hiring flexibility on employment elasticity is driven by the reduction in hiring frictions. Second, when there is lower job creation, firms have incentives to create more jobs. But this effect can be exactly (or more than) offset by increases in unemployment benefits that result in higher wages and thus reduce job match profitability.

A less-studied aspect of labour market policies is the impact of ALMPs such as training programmes. When training expenditure is assumed to be a mere cost to firms (as theoretically modelled by e.g. Onwordi, 2016; Pissarides, 2009; Stähler and Thomas, 2012), it leads to lower job match surplus and profitability to the firms. In this instance, even a small productivity shock can lead to large fluctuations in employment. However, as argued by Molana, Montagna and Onwordi (2018), training is likely to be productivity enhancing; allowing for this, they show that training expenditure can have a moderating effect on employment fluctuations in response to productivity shocks. We however fail to find empirical support for these predictions since our results show mostly statistically insignificant effects.

To better appreciate the impact of these policies, we study their complementarities more systematically. Particularly, we examine the effects of, ceteris paribus, a one-standard-deviation upward move in each policy variable under three labour market regimes: *average*, *liberal* and *interventionist*. The policy regimes differ in the values that the non-changing policy variables take. The hypothetical *average* regime sets the policy variables at their sample average values. Because the policy variables in the regressions are centered, their average value is zero. The non-centered averages are 0.309 for *grr*, 7.116 for *train* and 4.114 for *flex* for the sample that excludes Spain (Table 3 reports the non-centered averages with Spain). The *liberal* regime assumes a flexibility index at the 75th percentile of its sample distribution (*flex* = 4.955, non-centered) and the other two policy variables at their 25th percentiles (*grr* = 0.219, *train* = 6.626, both non-centered). The *interventionist* regime, in contrast, is characterised by a flexibility index at its 25th percentile (*flex* = 3.425) and the other two at their 75th percentiles (*grr* = 0.388, *train* = 7.919). These inter-quartile differences roughly coincide with the average differences between the typical liberal and interventionist countries in our sample.

We rely on expression (5) to calculate the effects on the short-run elasticity, using the estimated coefficients in Table A.11, reported either in column 4 (without policy interactions) or in column 5 (with policy interactions). Without policy interactions, the effect on the short-run elasticity of a one-standard-deviation upward move in the gross replacement rate, for instance, is obtained as $\hat{\gamma}_2^{\tilde{\text{grr}}} \times \hat{\sigma}^{\text{grr}} = 0.843 \times 0.122 = 0.103$, where $\hat{\gamma}_2^{\tilde{\text{grr}}}$ is the estimated coefficient for the interaction of Δy with the gross replacement rate in column 4, and $\hat{\sigma}^{\text{grr}}$ is the sample standard deviation of the gross replacement rate. With policy interactions, the other policy variables will also matter and the change in the short-run elasticity becomes

$$sre \text{ change} = \left(\hat{\gamma}_2^{\tilde{\text{grr}}} + \hat{\gamma}_2^{\tilde{\text{grr}} \text{ x } \tilde{\text{train}}} \times \tilde{\text{train}} + \hat{\gamma}_2^{\tilde{\text{grr}} \text{ x } \tilde{\text{flex}}} \times \tilde{\text{flex}} + \hat{\gamma}_2^{\tilde{\text{grr}} \text{ x } \tilde{\text{train}} \text{ x } \tilde{\text{flex}}} \times \tilde{\text{train}} \times \tilde{\text{flex}}\right) \times \hat{\sigma}^{\text{grr}},$$

where the $\hat{\gamma}_2$s are the coefficients for the corresponding policy interactions with Δy in column 5 of Table A.11 and the policy variables $\tilde{\text{train}}$ and $\tilde{\text{flex}}$ are set at their regime-specific 'centered' values. The results of this exercise are reported on the left-hand side of Table 4, together with standard errors which reflect parameter uncertainty. For robustness, the right-hand side of the table also reports the estimates based on the regressions with additional LMIs in columns 6 and 7 of Table A.11.

Table 4: Effects on the short-run elasticity of σ-increases in the policy variables

	baseline				with additional LMIs			
Policy	without	with interactions			without	with interactions		
variable	interactions	average	liberal	interventionist	interactions	average	liberal	interventionist
grr	0.103***	0.122***	0.039	0.156**	0.131***	0.138***	0.096	0.155**
	(0.040)	(0.043)	(0.052)	(0.063)	(0.044)	(0.051)	(0.077)	(0.061)
train	0.037	0.077	-0.001	0.041	0.040	0.039	0.161	0.019
	(0.032)	(0.050)	(0.166)	(0.046)	(0.031)	(0.049)	(0.191)	(0.048)
flex	0.108***	0.111***	0.186***	0.073	0.062*	0.063	0.122**	0.010
	(0.033)	(0.035)	(0.049)	(0.054)	(0.037)	(0.040)	(0.060)	(0.056)

Notes: The figures show the effects of a 1-standard-deviation increase in a policy variable on the short-run employment-output elasticity. Baseline estimates are based on columns 4 (without interactions) and 5 (with interactions), those with additional LMIs on column 6 and 7 of Table A.11. Average, liberal and interventionist regimes differ in the values the non-changing policy variables are assumed to take. Standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

When policy complementarities are not allowed for, the estimated effects do not depend on the policy regime (first column): a move towards more unemployment benefit generosity and a more flexible labour market both correspond to larger short-run employment-output elasticities (both roughly by 0.1). When, however, policy complementarities are taken into account, it matters considerably in which labour market regime a certain policy change takes effect. Specifically, our estimation reveals that in *liberal* regimes, making the unemployment benefit more generous does not result in a significantly higher short-run elasticity. In a similar vein, a reform towards more labour market flexibility does not lead to a significantly higher short-run employment responsiveness under an *interventionist* policy mix. Indeed, as noted previously, it is possible that the effects of the two policies can exactly offset each other, which may explain the non-statistically significant effects. In particular, as argued by, e.g., Hornstein, Krusell and Violante (2005) in a paper that theoretically assesses labour market fluctuations, the increase in the generosity of unemployment benefit needs to be sufficiently high in order to induce large fluctuations in employment. This would be the case especially if accompanied by greater hiring and firing flexibility. An important implication of these results is, therefore, that the impact of a specific policy on employment fluctuations depends on the initial policy mix and levels.

These findings are qualitatively robust to the inclusion of additional LMIs in the regression (right-hand side of Table 4) or to estimating the FD specification of the model (Table A.13), although these changes make the impact of a move towards more flexibility qualitatively smaller. In contrast, the results for the *flex* indicator are not robust to the inclusion of Spain in the sample, which is shown on the right-hand side of Table A.14 in the Appendix. This is consistent with employment in Spain being very responsive to output shocks, which is however at odds with the comparetively low value of *flex* for Spain. As said, this contradiction is due to the fact that in Spain there is a very high share of temporary work contracts, where employment protection is significantly lower than for regular contracts (Bentolila et al., 2010), a feature not captured by our empirical model. To account for this, we reproduce the estimation with additional LMIs (and their interactions with output growth), where the set of LMIs also includes a variable for the share of temporary work contracts (temp). These results are shown on the right-hand side of Table A.14. Because variable temp is observed only for two-thirds of the observations in our sample, the results are not directly comparable to our baseline estimates. It is however notable that our earlier significant positive coefficient for *flex* is restored once the varying importance of temporary contracts across countries is accounted for.

Further, using the estimates from the EC regressions, we also calculate the corresponding changes in the long-run elasticity, which we report for the baseline case in Table A.15.[20] The numbers suggest that the effects of these policies found on the short-run elasticity mostly fade out in the long run. Interestingly, the point estimates for training turn negative in the long run (though being not statistically significant), which could suggest that training has a lock-in effect, reflecting a greater propensity of firms to hold on to trained employees in response to shocks (Molana, Montagna and Onwordi, 2018).[21]

7 Reforming towards flexicurity

In light of the results obtained so far, in this section, we consider the effects of a flexicurity reform on the employment-output elasticity. A flexicurity regime is characterised by a combination of generous unemployment insurance and training programmes and a relatively flexible labour market. Our specific objective is to examine the effects, in each country, of 'reforming' its existing policy mix in 2010 towards the policy mix characterising the Danish system, a well-known example of a flexicurity regime (e.g. as discussed in Koster *et al.*, 2011; Räisänen *et al.*, 2012).

Denmark has one of the highest unemployment benefit replacement rates and training expenditures among the countries in our sample, while it scores close to the average in the flexibility index. Table A.16 in the Appendix displays the countries' policy distances from Denmark in 2010. Clearly, these

[20] This we do by evaluating expression (6) before and after the one-standard-deviation increase in a given policy variable and taking the difference. Corresponding standard errors are obtained with the Delta method. For the calculation 'without interactions' we assume that policy variables take their average values. Taking other values does not influence the results qualitatively.

[21] In a search and matching open economy model, Molana, Montagna and Onwordi (2018) show that increases in training expenditure can reduce employment volatility in response to exogenous shocks even in the presence of more generous unemployment insurance.

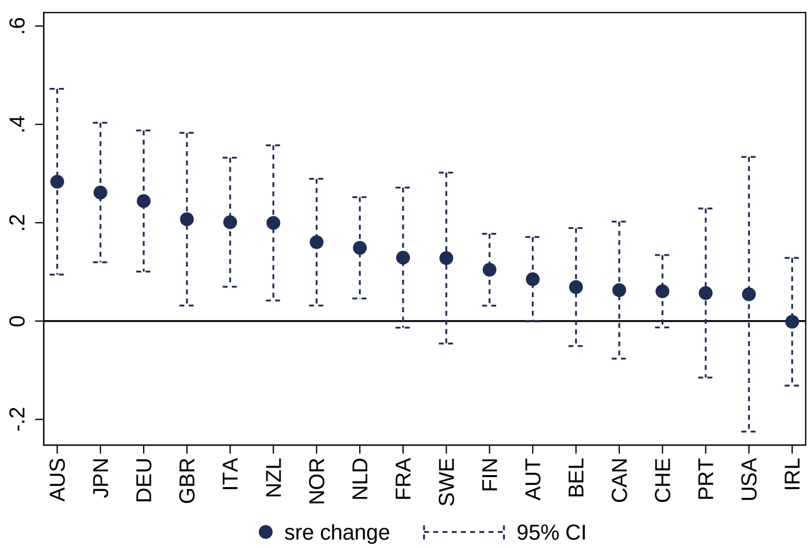
figures imply that a reform in the direction of flexicurity would require most countries within our sample to increase their benefit generosity and raise their expenditures on training, while, simultaneously adjusting the flexibility of their labour markets towards an average level. In light of this, our priors are that a reform in the direction of flexicurity is likely to increase or leave the short-run elasticity of all countries roughly unchanged. The main reason, as suggested by our data, is that this reform would make each country's regime more generous or leave them almost as generous as they were. Moreover, for most countries with regimes characterised by a generosity similar to Denmark, the reform would mean an increase in labour market flexibility. Less straightforward is the case of countries with *liberal* regimes, where the parallel moves towards less flexibility and more generosity work in opposite directions, as revealed in the previous section.

We report the estimated changes in the short-run elasticity resulting from the policy shift described above in Figure 4, in descending order, together with the 95% confidence intervals. The computation was based on expression (5) and the estimated coefficients in Table A.11, column 5, and performed as

$$sre\ \text{change}_i = \sum_k \hat{\gamma}_2^k \left[\tilde{P}^k_{\text{DNK},t=2010} - \tilde{P}^k_{i,t=2010}\right],$$

where $\tilde{P}^k$ denote the policy variables and their interactions and DNK stands for Denmark.

Figure 4: Impact of a flexicurity reform on the short-run elasticity



Countries in descending order of the change in short-run elasticity.

Consistent with our priors, we find that a flexicurity reform would increase or leave unchanged the short-run employment-output elasticity in all of our sample countries. This result is in line with the philosophy underlying the flexicurity system, which does not aim to prevent employment separations (hence the high short-run elasticity), but to provide targeted incentives and support for the unemployed to return to work. The highest increases of above 0.2 are obtained for Australia, Japan, Germany and Great Britain. Positive and significant effects are obtained also for Italy, New Zealand, Norway, the Netherlands and Finland. At the other end, the estimated effects are small or not statistically significant for, among others, Ireland, USA, Portugal and Canada.

The group of countries with the highest expected effects includes those with considerably less generous systems than Denmark (e.g. Australia, Great Britain). An initially less generous system, in itself, however does not guarantee a positive effect. For two of the least generous countries, the US and Canada, we do not find statistically significant effects. Under a flexicurity reform, these two countries would need to reduce the flexibility of their labour markets, which would counteract the effects of the reform package on output-employment dynamics.

The above implications of a flexicurity reform on the short-run elasticity are largely robust when relying on estimates from the FD specification, as can be seen by comparing the first columns of Table A.17 and Table A.18. The groups of countries for which statistically significant increases in the short-run elasticity are predicted are almost identical under the baseline and the FD specifications, and the relative sizes of the country estimates also remain largely unchanged. We obtain smaller estimates for most countries however when additional LMIs are included in the EC model. Nevertheless, the predicted effects for the four most affected countries (Australia, Japan, Germany, Great Britain) remain large and statistically significant even under this model.

These results indirectly lend further support to the importance of taking into account all – combined and complementary – effects when we evaluate a labour market reform.

As for the long-run elasticity, our model does not predict statistically significant changes (Figure A.6). Hence, no evidence is found that a flexicurity reform would have an impact on the responsiveness of employment to output in the long run.

8 Robustness

8.1 Heterogeneous country trends

Our simple model accounts for country differences that are constant over time but cannot account for the possibility that secular trends in the employment-output relationship due to, e.g., demography, productivity or technology, differ from country to country. In one robustness exercise, we aim to allow for such heterogeneous trends by adding country-specific linear time trends to the model. First, we reproduce Table 1 and Table A.7 while including country-specific linear time trends and find that the short-run elasticity estimates are robust to this change.[22] We also find evidence that the cointegrating relationship exists under the assumption of such trends. Nevertheless, for the short panel, the estimates for the speed of adjustment parameter and the long-run elasticity are significantly larger in absolute value than the baseline figures, and the long-run elasticity becomes larger than one and thus falls outside the conventional range of such estimates. This suggests that the inclusion of country trends in the EC regression on the short panel carries the risk of model overfitting. For this reason, we proceed with estimating the policy regressions only for the more parsimonious FD specification.

The left-hand side of Table 5 reports the estimated effects on the short-run elasticity of one-standard-error increases in the policy variables. It is based on regressions of the FD specification

[22]To save space, these estimation tables are not reported but are available upon request.

Table 5: Robustness with country trends and per-capita transformation

	country trends				per capita transformation			
Policy	without	with interactions			without	with interactions		
variable	interactions	average	liberal	interventionist	interactions	average	liberal	interventionist
grr	0.084**	0.077	0.049	0.124*	0.085*	0.091*	0.005	0.126*
	(0.041)	(0.047)	(0.056)	(0.067)	(0.046)	(0.048)	(0.054)	(0.066)
train	0.084**	0.069	0.156	0.077	0.044	0.064	0.056	-0.019
	(0.034)	(0.050)	(0.181)	(0.051)	(0.037)	(0.048)	(0.166)	(0.047)
flex	0.115***	0.085**	0.104**	0.051	0.077***	0.052	0.110**	0.045
	(0.034)	(0.036)	(0.051)	(0.059)	(0.030)	(0.032)	(0.045)	(0.052)

Notes: The figures show the effects of a 1-standard-deviation increase in a policy variable on the short-run employment-output elasticity based on regressions of the FD specification. The left part of the table reports results from regressions with linear country trends, the right part reports results from regressions with per-capita transformed employment and output. Additional LMIs are not included. The regressions with per-capita transformation exclude from the set of control variables the change in the working-age population growth. The sample excludes Spain. Average, liberal and interventionist regimes differ in the values the non-changing policy variables are assumed to take. Standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

without additional LMIs such as the left-hand side of Table A.13 but with the inclusion of country-specific linear trends. The inclusion of trends leaves the qualitative finding on the regime-specific effects unchanged. More benefit generosity increases the short-run elasticity only in relatively interventionist regimes, while a higher flexibility index raises the elasticity only in more liberal regimes. Finally, the simulated reform toward flexicurity using the model with country trends leads to very similar results to those without such trends, which is evident from comparing columns (1) and (2) in Table A.18.

8.2 Per capita measurement of employment and output

As a more direct test for the role of demographic forces, in one robustness check we estimate equation (2) with both employment and output divided by the size of the working age population (ages 20-64), a specification which directly accounts for changes in the size of the labour force. The per capita transformation, which essentially turns employment into employment rate and output into output per capita, largely retains the time series properties of the two variables. Reproducing Table 1 and Table A.7 shows that the estimates for the short-run elasticity are remarkably robust to the per-capita transformation of employment and output. Nevertheless, as opposed to the baseline estimates, the Ericsson-MacKinnon cointegration test cannot reject the null of no cointegration. This makes us reproduce the policy regressions with the FD specification, which does not assume the existence of cointegration. Our main results on the regime-specific effects of policy changes on the short-run elasticity are robust to the per-capita transformation (right-hand side of Table 5). Consequently, the derived effects of a flexicurity reform in individual countries remain similar to the baseline results (column (3) in Table A.18).

8.3 Broad measure of ALMP expenditures

In the main analysis, we represented the ALMP spendings of a country with its expenditures on ALMP-related training programmes. However, ALMPs also incorporate elements beyond training.

According to the OECD classification, the following categories of policies belong to ALMPs:[23] training (institutional, workplace, and integrated training, special support for apprenticeship); employment incentives (recruitment incentives, employment maintenance incentives, job rotation and job sharing); sheltered and supported employment and rehabilitation; direct job creation; and start-up incentives. As discussed earlier, our focus on training is motivated by the fact that training expenditures constitute the highest share of total spending on ALMPs and that training is clearly targetted at improving employability, which may be less true for other types of ALMPs. For instance, direct job creation such as public works programmes tends to be less effective in promoting employability, especially when they are not combined with training (e.g., Gehrke and Hartwig, 2018). This section tests the robustness of the results to using a broad ALMP measure, which combines all the categories listed above. Just as *train* in the main analysis, the broad ALMP variable is also expressed in PPP-adjusted US dollars and divided by the number of unemployed persons in the country. This broad ALMP has a sample mean of 7.52 with a standard deviation of 1.04 and it is very strongly correlated with *train* (correlation coefficient is 0.9).

Table 6: Robustness with broad ALMP

	without LMIs				with LMIs			
Policy	without	with interactions			without	with interactions		
variable	interactions	average	liberal	interventionist	interactions	average	liberal	interventionist
grr	0.118***	0.130***	0.066	0.099	0.140***	0.166***	0.115	0.149***
	(0.043)	(0.049)	(0.069)	(0.065)	(0.044)	(0.056)	(0.087)	(0.057)
almp	0.019	0.068	-0.163	0.016	0.011	0.060	-0.106	0.059
	(0.037)	(0.052)	(0.226)	(0.044)	(0.039)	(0.054)	(0.234)	(0.052)
flex	0.105***	0.098***	0.173***	0.103	0.066*	0.060	0.157***	0.017
	(0.033)	(0.035)	(0.053)	(0.063)	(0.039)	(0.039)	(0.058)	(0.059)

Notes: The figures show the effects of a 1-standard-deviation increase in a policy variable on the short-run employment-output elasticity. ALMPs are represented by a broad ALMP measure that also comprises training. The estimates are based on a regression of the EC specification either without or with additional LMIs. The sample excludes Spain. Average, liberal and interventionist regimes differ in the values the non-changing policy variables are assumed to take. Standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6 reports the estimated effects of σ-increases in the policy variables. The reported estimates are obtained from regressions identical to those behind Table 4 in all respects but the ALMP variable. We can conclude that the results are largely robust to this change. In particular, the estimates associated with the increase in ALMP spending remain statistically zero under all policy regimes. The flexicurity reform exercise also yields qualitatively similar results if we base it on the regressions with the alternative ALMP variable (column (3) in Table A.17).

8.4 Two-step estimation

Our empirical approach behind regression equation (2) is to estimate the long-run cointegrating relationship and the dynamic adjustment in a single step – an approach that has recently been proposed by Rao, Singh and Kumar (2010) and Kumar and Rao (2012). Alternatively, one could follow a two-step approach, first estimating the long-run cointegrating relationship between e_{it} and y_{it} and then,

[23] A detailed description of the categories can be found on the OECD's webpage here: `https://www.oecd.org/els/emp/Coverage-and-classification-of-OECD-data-2015.pdf`

given that there is cointegration, using the lagged first-stage residuals, i.e., the lagged error correction (ECM) term, in the second-stage regression for the dynamic relationship. An advantage of the two-step approach is that it allows for alternative estimation methods that have been proposed to estimate cointegrating relationships: the Fully Modified OLS (FMOLS) introduced by Phillips and Hansen (1990) and the Dynamic OLS (DOLS) proposed by Stock and Watson (1993). If the cointegrating equation error and the regressor innovations (in our case, innovations of output) are correlated, OLS estimates of the long-run cointegrating relationship can be substantially biased in small samples (albeit being super consistent) and the standard error estimates become invalid. To soak up the possibly problematic correlation, the FMOLS uses a semiparametric correction while the DOLS augments the cointegrating regression with leads and lags of the first-differenced regressors. These two methods produce asymptotically unbiased and normally distributed coefficient estimates.

In this robustness check we carry out two-step estimations, using either the FMOLS or the DOLS estimators in the first stage. Our aim is to reproduce the estimation results reported in Table 1 for the long panel and to show that the resulting estimates for the employment-to-output elasticity are close to the baseline estimates. In the first step, we estimate the cointegrating relationship between log employment (number of employees) and log output with FMOLS and DOLS, where we include as deterministic terms either a constant only or a constant and a linear time trend.[24] The upper part of Table A.19 presents the corresponding estimates for the long-run relationship between employment and output. Interestingly, the estimate is significantly smaller than our implied long-run elasticity from the baseline estimation, suggesting that the baseline estimate is upward biased. The presence of cointegration can be tested by checking whether the residuals obtained from these first-stage regressions are stationary. Conventional panel unit root tests with a null hypothesis of non-stationarity (Levin-Lin-Chu, Fisher-ADF) reject the null for all countries, while the Hadri tests with a null hypothesis of stationarity for *all* countries also reject the null. This is consistent with the results of the Johansen test performed earlier, where we found that cointegration exists for most - but not for all - countries in our sample. In the second step, we estimate the dynamic relationship in a regression similar to (2) but replace the variables in levels, $e_{i,t-1}$ and $y_{i,t-1}$, with the lagged ECM term. These estimates are reported in the lower part of Table A.19. Importantly, the short-run elasticity estimates are remarkably robust and remain at the level of the baseline estimate (0.3) regardless of the estimation method used in the first stage. Furthermore, the coefficient for the lagged residual, which corresponds to the speed of adjustment, is quantitatively very close to our baseline estimate and highly significant statistically.

9 Conclusion

The variations in employment responses to external shock and differences in its recovery pattern seen in the aftermath of the Great Recession are partly blamed on the nature of labour market policies and institutions characterising the individual countries within both the EU and the OECD. Consequently, the need to enhance the ability of the labour market to adjust to shocks has heightened the emphasis on *flexicurity* reforms. This paper examines the responsiveness of employment to output fluctuations

[24]Contrary to the rest of the analysis, this estimation is done in EViews and not in Stata because EViews has better functionalities for FMOLS and DOLS on panel data. DOLS includes zero or one lead and lag, depending on the Akaike IC. Default options are applied for the assumptions concerning the long-run variance.

in a panel of 20 OECD countries, paying particular attention to the role of the key pillars of flexicurity as well as their interactive effects in driving this outcome.

Interesting results emerge. For an average country, we find that there has been a significant increase in the responsiveness of employment to output fluctuations over the past decades, with labour market policies playing a crucial role in shaping this outcome. In particular, our results show that important complementarities exist between these policies, especially in determining the short-run employment-output dynamics, and that – depending on the initial policy mix – reforms in the direction of flexicurity may result in asymmetric outcomes for the individual economies. A clear conclusion that emerges from this study is that a one-size-fits-all policy may not yield the same results everywhere. Instead, in considering which reform to undertake, an economy should first consider its existing policy framework.

Of course, there may be other factors that remain unexplained in our estimation that may in fact be relevant in driving employment dynamics – for instance the changes in the nature of job creation in different countries (e.g. different reliance on flexible short-term contracts) – which we leave for future research.

References

Abbritti, M. and S. Fahr (2013) "Downward wage rigidity and business cycle asymmetries", J*ournal of Monetary Economics*, 60(7), 871-886.

Abbritti, M. and S. Weber (2010) "Labor market institutions and the business cycle – unemployment rigidities vs real wage rigidities", ECB WP series No. 1183.

Abbritti, M. and S. Weber (2018) "Reassessing the role of labor market institutions for the business cycle", *52nd issue of the International Journal of Central Banking.*

Andersen, T. M. (2015), "A Flexicurity Labor Market During Recession", *IZA World of Labor*, 173, 1–10.

Andersen, T. M., and M. Svarer (2007), "Flexicurity: Labour market performance in Denmark", CESifo Economic Studies, 53, 389–429.

Azariadis, C., and C. A. Pissarides (2007), "Unemployment dynamics with international capital mobility", *European Economic Review*, 51(1), 27-48.

Bassanini, A. (2012), "Aggregate Earnings and Macroeconomic Shocks: The Role of Labour Market Policies and Institutions", *Review of Economics and Institutions*, 3, 1–44.

Baxter, M., and R. G. King (1999), "Measuring business cycles: Approximate band-pass filters for economic time series", *Review of Economics and Statistic*s, 81, 575-593.

Behar, A. (2017), "Comparing the employment-output elasticities of migrants and nationals: the case of the Gulf Cooperation Council", *Applied Economics Letters*, 24(10), 703-707.

Behar, A. and J. Hodge (2008), "The Employment Effects of Mergers in a Declining Industry: The Case of South African Gold Mining", *The B.E. Journal of Economic Analysis & Policy*, 8.

Bertola, G., F. D. Blau and L. M. Kahn (2002), "Comparative Analysis of Labor Market Outcomes: Lessons for the US from International Long-Run Evidence", In: A.B. Krueger and R. Solow (eds.) *The Roaring Nineties: Can Full Employment Be Sustained?*, Russell Sage Foundation and The Century Foundations, New York. 159–218.

Bentolila, S., P. Cahuc, J. J. Dolado, and T. Le Barbanchon (2010) "Unemployment and Temporary Jobs in the Crisis: Comparing France and Spain", Working Papers 2010-07, FEDEA.

Blanchard, O., C. Bean, and W. Münchau (2006), "European Unemployment: The Evolution of Facts and Ideas", *Economic Policy*, 21(45), 7–59.

Blanchard, O.J., Jaumotte, F. and P. Loungani, (2014). "Labor market policies and IMF advice in advanced economies during the Great Recession", *IZA Journal of Labor Policy*, 3, 1–23.

Blanchard, O. and J. Wolfers (2000), "The Role of Shocks and Institutions in the Rise of European Unemployment: The Aggregate Evidence", *Economic Journal*, 110, C1–C33.

Bowdler, C. and L. Nunziata (2007), "Inflation Adjustment and Labour Market Structures: Evidence from a Multi-country Study", *Scandinavian Journal of Economics*, 109, 619–642.

Cacciatore, M. (2014), "International Trade and Macroeconomic Dynamics with Labour Market Frictions", *Journal of International Economics*, 93, 17-30.

Cacciatore, M., G. Fiori and F. Ghironi (2016), "Market Deregulation and Optimal Monetary Policy in a Monetary Union", *Journal of International Economics*, 99, 120–137.

Card, D., J. Kluve and A. Weber (2010) "Active Labour Market Policy Evaluations: A Meta-Analysis", *Economic Journal*, 120, F452–F477.

Ciminelli, G., R. Duval, and D. Furceri (2020), "Employment Protection Deregulation and Labor Shares in Advanced Economies", *The Review of Economics and Statistics*, 1-44.

Crivelli, E., D. Furceri and J. Toujas-Bernate (2012), "Can Policies Affect Employment Intensity of Growth? A Cross-Country Analysis", IMF Working Paper No. 12/218.

Dabusinskas, A., I. Kónya, and S. Millard (2016), "How does labour market structure affect the response of economies to shocks?" Staff Working Paper No. 582. Bank of England.

Davoine, T., and C. Keuschnigg (2015), "Flexicurity, Taxes and Job Reallocation", CESIfo Working Paper 5302.

Duval, R., D. Furceri, and J. Jalles (2020), "Job protection deregulation in good and bad times", *Oxford Economic Papers*, 72(2), 370–390.

Eichhorst, W. and R. Konle-Seidl (2006), "The Interaction of Labor Market Regulation and Labor Market Policies in Welfare State Reform", *Comparative Labor Law & Policy Journal*, 28, 1.

Ericsson, N. R. and J. G. MacKinnon (2002), "Distributions of Error Correction Tests for Cointegration" *Econometrics Journal*, 5(2), 285–318.

Faccini, R. and C. R. Bondibene (2012), "Labour market institutions and unemployment volatility: evidence from OECD countries", Bank of England, Working Paper, No. 461.

Feenstra, R. C., R. Inklaar and M. P. Timmer (2015), "The Next Generation of the Penn World Table", *American Economic Review*, 105(10), 3150–3182.

Fonseca, R., L. Patureau, and T. Sopraseuth (2010), "Business Cycle Comovement and Labor Market Institutions: An Empirical Investigation", *Review of International Economics*, 18, 865–881.

Gehrke, E. and R. Hartwig (2018) "Productive effects of public works programs: What do we know? What should we know?", *World Development*, 107, 111-124.

Görg, H., Henze, P., Jienwatcharamongkhol, V., Kopasker, D., Molana, H., Montagna, C., and Sjöholm, F. (2017), "Firm size distribution and employment fluctuations: Theory and evidence", *Research in Economics*, 71(4), 690-703.

Gnocchi S., A. Lagerborg, and E. Pappa (2015), "Do labor market institutions matter for business cycles?", *Journal of Economic Dynamics & Control*, 51, 299–317.

Hagedorn, M. and I. Manovskii (2008), "The Cyclical Behavior of Equilibrium Unemployment and Vacancies Revisited", *American Economic Review*, 98, 1692–1706.

Harding, D. and A. Pagan (2002). "Dissecting the cycle: A methodological investigation", *Journal of Monetary Economics*, 49, 365–81.

Hendry, D. F. (1995), *Dynamic Econometrics*, Oxford University Press, Oxford, United Kingdom.

Hijzen, A. and S. Martin (2013) "The role of short-time-work schemes during the Global Financial Crisis and Early Recovery: A Cross-Country Analysis", *IZA Journal of Labor Policy*, 2, 5.

Hornstein, A., P. Krusell, and G. L. Violante (2005), "Unemployment and Vacancy Fluctuations in the Matching Model: Inspecting the Mechanisms", *Federal Reserve Bank of Richmond Economic Quarterly*, 91, 19–51.

Johansen, S. (1991), "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models", *Econometrica*, 59, 1551–1580.

Kapsos, S. (2006), "The Employment Intensity of Growth: Trends and Macroeconomic Determinants", In: Felipe J., Hasan R. (eds) *Labor Markets in Asia.* Palgrave Macmillan, London

Kluve, J. (2010), "The Effectiveness of European Active Labor Market Programs", *Labour Economics*, 17, 904–918.

Kolasa, M., M. Rubaszek, and M. Walerych (2021) "Do flexible working hours amplify or stabilize unemployment fluctuations", *European Economic Reviw*, 131, 103605.

Koster, F., J. Mcquinn, I. Siedschlag, and O. van Vliet (2011), "Labour Market Models in the EU", NEUJOBS Special Report, No. 1.

Kumar, S. and B. B. Rao (2012) “Error-correction based panel estimates of the demand for money of selected Asian countries with the extreme bounds analysis”, *Economic Modelling*, 29, 1181–1188.

MacKinnon, J. G. (2010), “Critical values for cointegration tests”, Queen’s Economics Department Working Paper 1227.

McKay, A. and R. Reis (2008) “The brevity and violence of contractions and expansions”, *Journal of Monetary Economics*, 55, 738– 751.

Molana, H., C. Montagna, G.E. Onwordi (2018), “Reforming the Liberal Welfare State: Flexicurity and Labour Market Outcomes”, University of Aberdeen.

Molana, H., C. Montagna, G.E. Onwordi (2021), “De-Globalisation, Welfare State Reforms and Labour Market Outcomes”, *IMF Economic Review*, 69, 624–655.

Nickell, S., L. Nunziata, and W. Ochel (2005), “Unemployment in the OECD since the 1960s. What Do We Know?” *The Economic Journal*, 115(500), 1–27.

OECD (1994), *The OECD Jobs Study: Facts, Analysis, Strategies*, OECD, 1994.

OECD (2013), “Protecting jobs, enhancing flexibility: A new look at employment protection legislation”, Chapter 2 of the 2013 Edition of the OECD Employment Outlook.

Onwordi, G. E. (2016), “Labour Market Policies and Unemployment in the Presence of Search & Matching Frictions”, University of Dundee, PURE.

Pedroni, P. (1999), “Critical Values for Cointegration Tests in Heterogeneous Panels with Multiple Regressors”, *Oxford Bulletin of Economics and Statistics*, 61, 653-70.

Phillips, P. C. B. and B. E. Hansen (1990), “Statistical Inference in Instrumental Variables Regression with I(1) Processes”, *The Review of Economic Studies*, 57(1), 99–125.

Pissarides, C. A. (2009), “The Unemployment Volatility Puzzle: Is Wage Stickiness the Answer?”, *Econometrica*, 77, 1339–1369.

Rao, B. B., R. Singh and S. Kumar (2010) “Do we need time series econometrics?”, *Applied Economics Letters*, 17(7), 695-697.

Räisänen, H., J. Alatalo, K.K. Henriksen, T. Israelsson and S. Klinger (2012), “Labour Market Reforms and Performance in Denmark, Germany, Sweden and Finland”, Employment and Entrepreneurship, Ministry of Employment and the Economy, Helsinki.

Rumler, F and J. Scharler (2009), “Labor market institutions and macroeconomic volatility in a panel of OECD countries”, *Scottish Journal of Political Economy*, 583, 396–413.

Sahin, A., A. Tansel and M. H. Berument (2015), “Output-Employment Relationship Across Sectors: A Long- Versus Short-run Perspective,” *Bulletin of Economic Research*, 67(3), 265–288.

Shahidi, F. V., D. De Moortel, C. Muntaner, O. Davis and A. Siddiqi (2016), “Do Flexicurity Policies Protect Workers from the Adverse Health Consequences of Temporary Employment? A cross-national comparative analysis”, *SSM-Population Health*, 2, 674–682.

Shimer, R. (2005), “The Cyclical Behavior of Equilibrium Unemployment and Vacancies”, *American Economic Review*, 95, 25–49.

Stähler, N. and C. Thomas (2012), “FiMod – A DSGE Model for Fiscal Policy Simulations”, *Economic Modelling*, 29, 239–261.

Stock, J. H. and M. W. Watson (1993), “A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems”, *Econometrica*, 61(4), 783-820.

Westerlund, J. (2007), “Testing for Error Correction in Panel Data”, *Oxford Bulletin of Economics and Statistics*, 69, 709–748.

Zanetti, F. (2011), “Labor Market Institutions and Aggregate Fluctuations in a Search and Matching Model”, *European Economic Review*, 55, 644–658

Online Appendix

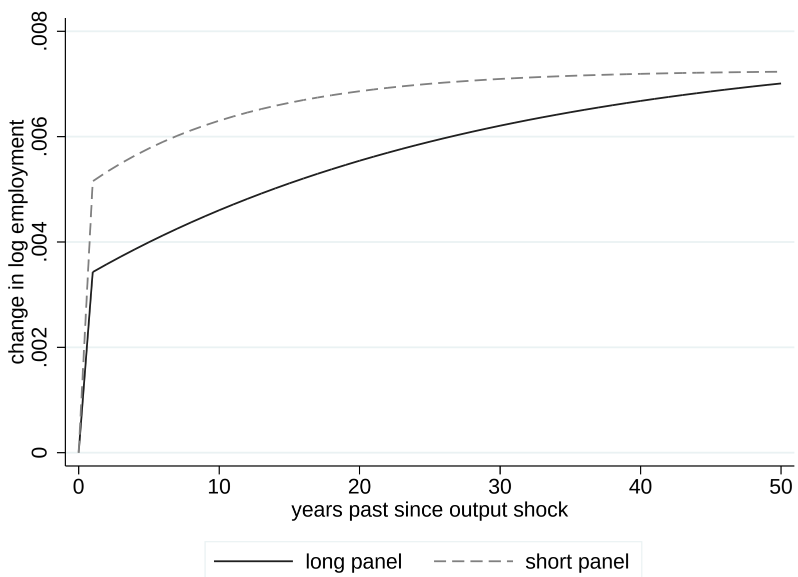
“Employment to Output Elasticities & Reforms towards Flexicurity: Evidence from OECD Countries”

Holger Görg* Cecília Hornok† Catia Montagna‡ George E. Onwordi§

September 16, 2022

Figures and Tables

Figure A.1: Employment response to a 1% permanent output shock



*Kiel Centre for Globalization, Kiel Institute for the World Economy and University of Kiel, Germany. E-mail: Holger.Goerg@ifw-kiel.de

†Kiel Centre for Globalization and Kiel Institute for the World Economy, Germany. E-mail: Cecilia.Hornok@ifw-kiel.de

‡University of Aberdeen and the Scottish Institute for Research in Economics, Scotland, UK. E-mail: c.montagna@abdn.ac.uk

§University of Aberdeen, Scotland, UK. E-mail: george.onwordi@abdn.ac.uk

Figure A.2: GDP growth cycle and years of slowdown by country

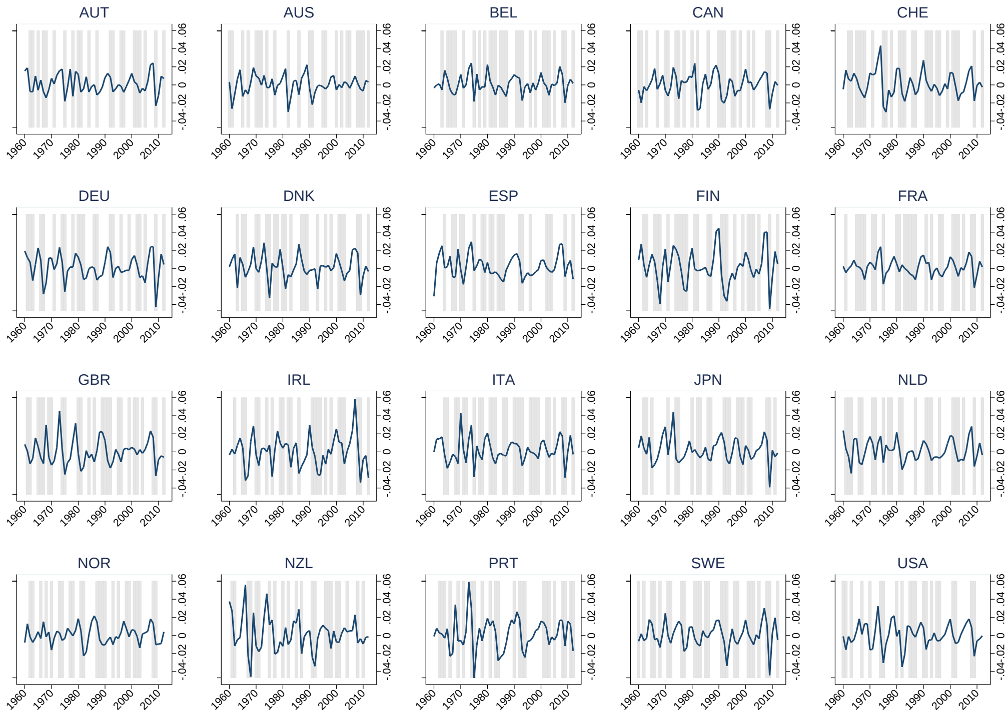


Figure A.3: Benefit replacement rate by country

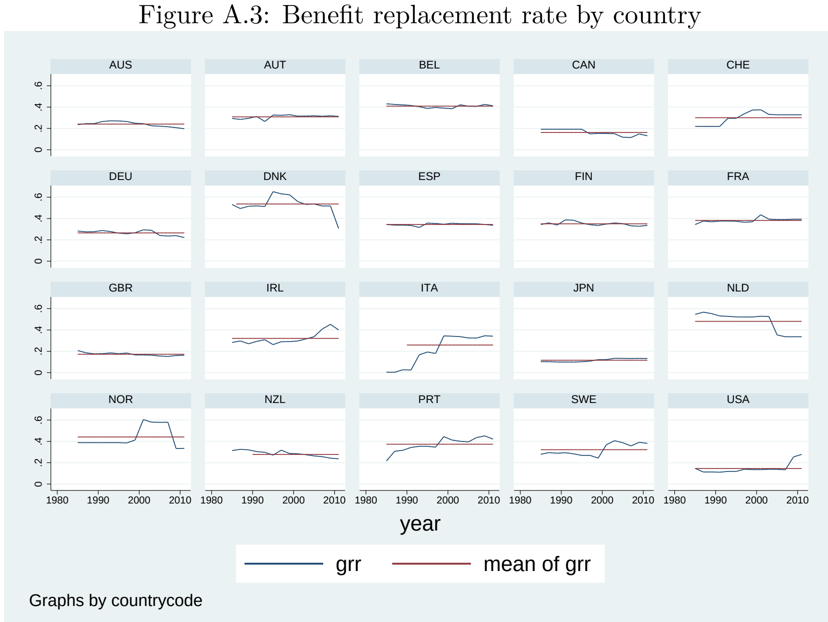


Figure A.4: Labour market flexibility by country

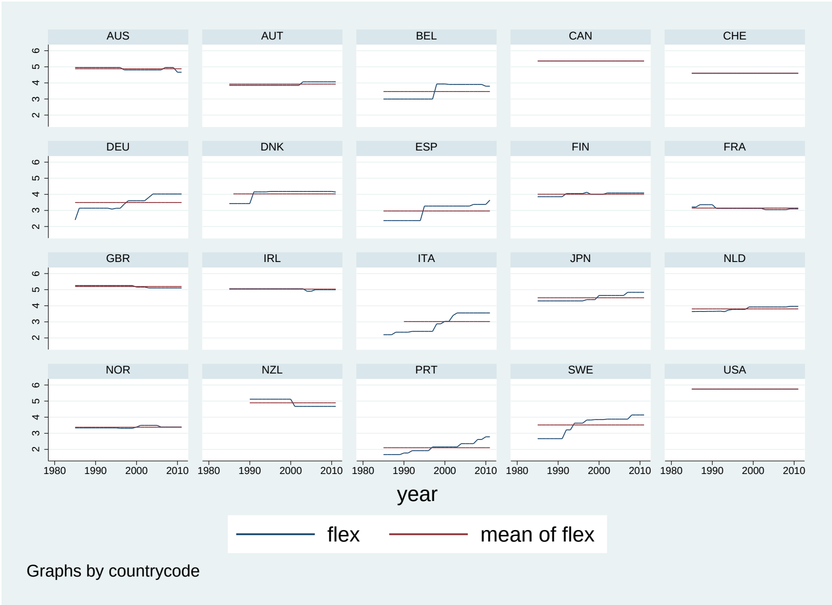


Figure A.5: Training expenditures per unemployed by country

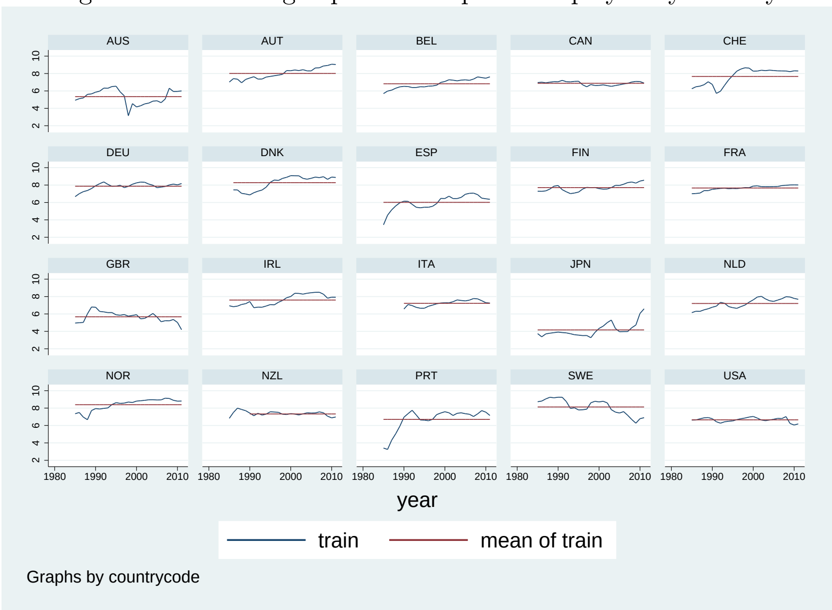


Table A.1: Description of variables used

Variable	Description	Data source
	Core variables	
employment	Number of persons engaged (millions)	Penn World Table 10.0
output	Real GDP at constant national prices (mil. 2017USD)	Penn World Table 10.0
avh	Average annual hours worked by person engaged	Penn World Table 10.0
	Policy variables	
grr	Gross unemployment benefit replacement rate	OECD
train	Training expenditures per unemployed person (PPP USD)	OECD and own calculation
flex	Labour market flexibility indicator (index, 0-6 scale)	OECD, inverse of EPL indicator
uden	Trade union density (share)	OECD
taxw	Tax wedge (share)	Abbritti and Weber (2018)
cove	Adjusted bargaining/union coverage rate (share)	OECD/AIAS ICTWSS database
cent	Centralisation of collective bargaining (index, 0-5 scale)	OECD/AIAS ICTWSS database
part	Prevalence of part-time employment (1-avh/fulltime hours)	PWT10.0 and own calculation
temp	Share of temporary employed in employed	OECD
	Control variables	
wapop	Working-age population (ages 20-64, mil. persons)	OECD
openness	Trade openness (export+import)/GDP in real 2010 USD	World Bank, WDI, and own calculation
govsize	Government consumption as share of GDP	OECD, National Accounts

Notes: The Penn World Table 10.0 is described in Feenstra, Inklaar and Timmer (2015).

Table A.2: ADF and KPSS unit root test statistics for levels

Country	log employment (e)					log output (y)				
code (ISO)	ADF	KPSS0	KPSS1	KPSS2	KPSS3	ADF	KPSS0	KPSS1	KPSS2	KPSS3
AUS	-1.992	0.313	0.168	0.122	0.101	-2.474	0.476	0.256	0.185	0.150
AUT	-0.769	0.937	0.499	0.354	0.283	-1.442	0.985	0.522	0.364	0.286
BEL	-0.263	1.073	0.554	0.381	0.296	-2.503	0.944	0.503	0.354	0.279
CAN	-1.293	1.006	0.532	0.377	0.301	-2.090	0.929	0.496	0.351	0.280
CHE	-1.584	0.354	0.190	0.138	0.113	-3.870	0.437	0.243	0.178	0.147
DEU	-0.899	0.334	0.185	0.139	0.119	-1.772	1.037	0.555	0.392	0.310
DNK	-2.418	0.410	0.231	0.175	0.150	-2.293	0.755	0.417	0.302	0.245
ESP	-0.986	1.033	0.525	0.359	0.278	-3.755	0.702	0.387	0.279	0.226
FIN	-1.221	0.366	0.192	0.136	0.110	-1.458	0.750	0.408	0.293	0.237
FRA	-1.606	0.605	0.318	0.225	0.181	-2.627	1.057	0.561	0.394	0.310
GBR	-0.940	1.016	0.530	0.372	0.295	-1.855	0.184	0.105	0.080	0.069
IRL	-1.427	1.059	0.543	0.372	0.288	-0.837	0.540	0.283	0.198	0.157
ITA	-2.514	0.470	0.251	0.181	0.147	-0.797	1.173	0.623	0.435	0.340
JPN	-0.413	1.124	0.588	0.408	0.319	-2.524	1.176	0.619	0.431	0.337
NLD	-1.452	0.603	0.320	0.228	0.183	-1.359	0.706	0.377	0.267	0.213
NOR	-1.728	0.339	0.180	0.131	0.109	0.093	1.053	0.561	0.396	0.315
NZL	-1.688	0.512	0.268	0.188	0.149	-2.249	0.469	0.254	0.183	0.148
PRT	1.177	0.629	0.348	0.252	0.204	-0.172	1.008	0.543	0.384	0.305
SWE	-1.635	0.686	0.359	0.254	0.202	-3.120	0.496	0.272	0.196	0.159
USA	0.090	1.043	0.554	0.394	0.316	-1.525	0.465	0.261	0.197	0.168

Notes: Annual series 1960-2012. ADF: Augmented Dickey-Fuller test. The test regression is specified to include a constant and a time trend and no lags. H0: series has a unit root (against being trend stationary). ADF critical values: 1% -4.146, 5% -3.498, 10% -3.179. KPSS: Kwiatkowski, Phillips, Schmidt, Shin test. H0: series is trend stationary. The maximum lag order is 3 for all countries; the test statistics for all lags from 0 to 3 are reported. KPSS critical values: 10% 0.119, 5% 0.146, 1% 0.216.

Table A.3: ADF and KPSS unit root test statistics for first differences

Country	dlog employment (Δe)					dlog output (Δy)				
code (ISO)	ADF	KPSS0	KPSS1	KPSS2	KPSS3	ADF	KPSS0	KPSS1	KPSS2	KPSS3
AUS	-4.557	0.110	0.078	0.072	0.072	-5.627	0.135	0.116	0.111	0.110
AUT	-3.405	0.077	0.053	0.047	0.044	-5.290	0.091	0.085	0.088	0.082
BEL	-4.425	0.201	0.144	0.125	0.118	-5.237	0.157	0.148	0.138	0.131
CAN	-4.849	0.050	0.038	0.037	0.041	-4.733	0.105	0.088	0.087	0.090
CHE	-3.962	0.151	0.096	0.081	0.076	-4.763	0.208	0.149	0.134	0.127
DEU	-3.987	0.107	0.072	0.065	0.066	-5.229	0.044	0.039	0.047	0.057
DNK	-4.512	0.085	0.063	0.059	0.061	-5.415	0.104	0.096	0.098	0.098
ESP	-2.223	0.260	0.152	0.116	0.098	-3.058	0.359	0.230	0.183	0.155
FIN	-3.327	0.138	0.084	0.070	0.066	-4.593	0.065	0.049	0.047	0.048
FRA	-4.205	0.089	0.061	0.055	0.054	-3.359	0.200	0.155	0.143	0.136
GBR	-3.661	0.077	0.050	0.045	0.046	-5.482	0.075	0.062	0.063	0.069
IRL	-3.442	0.254	0.160	0.125	0.110	-3.849	0.250	0.166	0.137	0.119
ITA	-3.872	0.147	0.098	0.085	0.078	-4.471	0.047	0.049	0.057	0.057
JPN	-3.283	0.097	0.067	0.057	0.052	-3.253	0.153	0.124	0.112	0.103
NLD	-4.362	0.127	0.089	0.081	0.085	-3.838	0.163	0.123	0.111	0.103
NOR	-3.311	0.092	0.057	0.049	0.048	-4.159	0.070	0.053	0.049	0.051
NZL	-4.554	0.169	0.121	0.106	0.102	-6.157	0.112	0.099	0.095	0.093
PRT	-3.799	0.379	0.272	0.235	0.208	-3.897	0.073	0.059	0.059	0.058
SWE	-3.871	0.128	0.084	0.077	0.079	-5.054	0.174	0.137	0.136	0.133
USA	-4.494	0.082	0.062	0.062	0.071	-4.977	0.074	0.060	0.060	0.066

Notes: First differences of annual series 1960-2012. ADF: Augmented Dickey-Fuller test. The test regression is specified to include a constant and 1 lag of the dependent variable. H0: series has a unit root. ADF critical values:1% -3.579 5% -2.929 10% -2.600. KPSS: Kwiatkowski, Phillips, Schmidt, Shin test. H0: series is trend stationary. The maximum lag order is 3 for all countries; the test statistics for all lags from 0 to 3 are reported. KPSS critical values: 10% 0.119, 5% 0.146, 1% 0.216.

Table A.4: Panel unit root tests for levels

	log employment (e)		log output (y)	
	Statistic	p-value	statistic	p-value
Levin-Lin-Chu				
Ho: Panels contain unit roots; Ha: Panels are stationary				
Adjusted t*	-2.089	0.018	-1.508	0.066
Fisher-ADF				
Ho: All panels contain unit roots; Ha: At least one panel is stationary				
Inverse chi-squared(40) P	38.091	0.557	39.435	0.496
Inverse normal Z	1.220	0.889	1.079	0.860
Inverse logit t(104) L*	1.407	0.919	1.422	0.921
Modified inv. chi-squared Pm	-0.214	0.585	-0.063	0.525
Hadri				
Ho: All panels are stationary; Ha: Some panels contain unit roots				
z	65.789	0.000	67.233	0.000

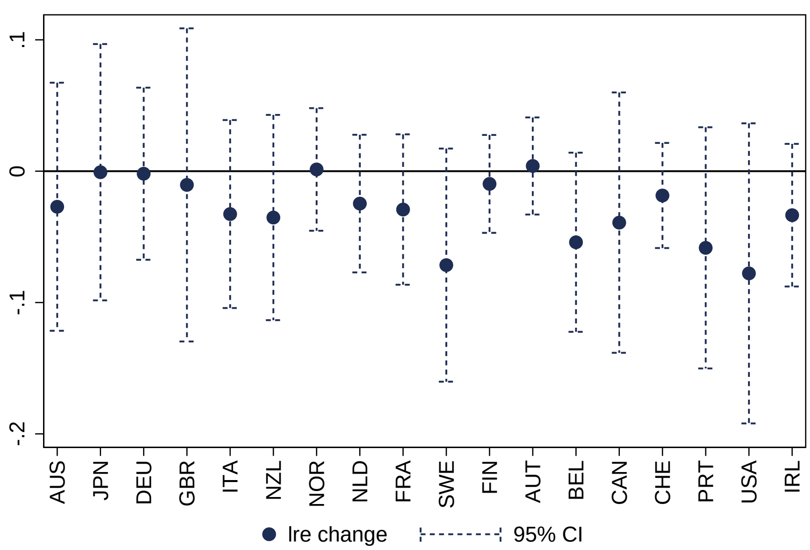
Notes: Annual panel of 20 countries over 1960-2012. For all tests country means are removed and a time trend is included. Fisher-ADF is performed with 3 lags.

Table A.5: Panel unit root tests for first differences

	dlog employment (Δe)		dlog output (Δy)	
	Statistic	p-value	statistic	p-value
Levin-Lin-Chu				
Ho: Panels contain unit roots; Ha: Panels are stationary				
Adjusted t*	-9.652	0.000	-11.956	0.000
Fisher-ADF				
Ho: All panels contain unit roots; Ha: At least one panel is stationary				
Inverse chi-squared(40) P	161.486	0.000	178.633	0.000
Inverse normal Z	-8.795	0.000	-9.658	0.000
Inverse logit t(104) L*	-9.802	0.000	-10.903	0.000
Modified inv. chi-squared Pm	13.583	0.000	15.500	0.000
Hadri				
Ho: All panels are stationary; Ha: Some panels contain unit roots				
z	10.456	0.000	12.715	0.000

Notes: Annual panel of 20 countries over 1960-2012. For all tests country means are removed. Fisher-ADF is performed with 2 lags.

Figure A.6: Impact of a flexicurity reform on the long-run elasticity



Countries in descending order of the change in short-run elasticity.

Table A.6: Johansen rank test for cointegration

Country code (ISO)	Max rank	Trace statistics	Max-eigenvalue statistics	Number of CI[1]	Number of CI[1]
AUS	0	25.073	17.066	0	1
	1	8.007	8.007		
AUT	0	40.147	31.551	1	1
	1	8.596	8.596		
BEL	0	51.410	41.372	1	1
	1	10.037	10.037		
CAN	0	28.149	20.116	1	1
	1	8.033	8.033		
CHE	0	35.932	25.117	1	1
	1	10.814	10.814		
DEU	0	30.457	20.732	1	1
	1	9.725	9.725		
DNK	0	35.495	26.192	1	1
	1	9.304	9.304		
ESP	0	43.164	34.738	1	1
	1	8.427	8.427		
FIN	0	31.256	24.051	1	1
	1	7.205	7.205		
FRA	0	38.834	30.128	1	1
	1	8.707	8.707		
GBR	0	18.394	11.094	0	1
	1	7.300	7.300		
IRL	0	16.089	11.679	0	1
	1	4.410	4.410		
ITA	0	44.003	30.817	1	1
	1	13.186	13.186		
JPN	0	28.127	22.795	1	1
	1	5.332	5.332		
NLD	0	35.519	27.917	1	1
	1	7.603	7.603		
NOR	0	36.701	25.924	1	1
	1	10.776	10.776		
NZL	0	18.072	11.913	0	1
	1	6.159	6.159		
PRT	0	30.263	19.063	1	1
	1	11.200	11.200		
SWE	0	26.107	20.609	1	1
	1	5.499	5.499		
USA	0	13.052	12.017	0	1
	1	1.035	1.035		
5% ciritical value	0	25.32	18.96		
	1	12.25	12.52		

Notes: Results from Johansen's cointegrating rank test for employment (number of employees) and output, 1960-2012, performed with restricted trend and 2 lags in the VAR. [1]Number of cointegrating equations chosen by multiple trace tests with 5% siginificance level. [2]Number of cointegrating equations chosen by minimizing HQIC (Hannan and Quinn Information Criterion).

Table A.7: Elasticity estimates (Spain excluded)

Depvar: Δe	Number of employees				FTE employment			
	1960-2012		1986-2012		1960-2012		1986-2012	
Δy (SR elasticity)	0.303***	0.324***	0.486***	0.483***	0.392***	0.408***	0.553***	0.564***
	(0.029)	(0.028)	(0.045)	(0.041)	(0.032)	(0.031)	(0.054)	(0.050)
e_{-1} (Speed of adj)		-0.033***		-0.089***		-0.034***		-0.092***
		(0.004)		(0.014)		(0.004)		(0.014)
		[-8.13]		[-6.16]		[-8.37]		[-6.46]
y_{-1}		0.026***		0.064***		0.018***		0.050***
		(0.005)		(0.012)		(0.006)		(0.015)
Observations	988	988	513	513	976	976	513	513
R-squared	0.480	0.521	0.613	0.649	0.500	0.529	0.617	0.651
LR elasticity		0.779***		0.719***		0.543***		0.546***
		(0.141)		(0.099)		(0.161)		(0.125)
CI test 5% crit val		-3.97		-4.15		-3.97		-4.15

Notes: Estimation of models (2) and (3) are based on OLS, where employment is either the number of employees or fulltime equivalent (FTE) employment. Country and year dummies are included in all regressions. Parentheses include robust standard errors, brackets t-statistics. Standard errors for the long-run elasticity are obtained with the delta method. The CI test is the Ericsson-MacKinnon cointegration test, with null hypothesis of no cointegration. Critical values (obtained from MacKinnon, 2010) must be compared with the t-statistic for the speed of adjustment parameter estimate. We can reject the null of no cointegration. *** p<0.01, ** p<0.05, * p<0.1.

Table A.8: Observations of economic slowdown

	All countries		European countries	
Period	Number	Share	Number	Share
1960-1969	88	0.440	69	0.460
1970-1979	86	0.430	66	0.440
1980-1989	91	0.455	72	0.480
1990-1999	98	0.490	72	0.480
2000-2012*	118	0.454	91	0.467
All (1960-2012)	481	0.454	370	0.465

Notes: Slowdown years are years of the contractionary phase of the growth cycle. The cycle is computed using band-pass filter and peaks and troughs are selected based on Harding and Pagan (2002) but using annual series. *Periods are decades except for the last period that includes 13 years.

Table A.9: Cross-country correlation of policies

	(grr,train)	(grr,flex)	(flex,train)
Full period			
1985-2011	0.595***	-0.527**	-0.255
Sub-periods			
1985-1994	0.322	-0.289	-0.071
1995-2004	0.634***	-0.552**	-0.283
2005-2011	0.727***	-0.693***	-0.415*

Notes: Correlation coefficients of period means. N=20. * significant at 10%, ** at 5%, *** at 1%.

Table A.10: Fisher-ADF panel unit root tests for control variables

	Level		First diff		Second diff	
	Statistic	p-value	Statistic	p-value	Statistic	p-value
			Variable: wapop			
Inverse chi-squared(40) P	35.484	0.674	48.195	0.175	261.060	0.000
Inverse normal Z	1.898	0.971	-0.714	0.238	-12.648	0.000
			Variable: openness			
Inverse chi-squared(40) P	50.932	0.1153	115.186	0.000		
Inverse normal Z	-0.911	0.181	-6.578	0.000		
			Variable: govsize			
Inverse chi-squared(40) P	47.903	0.183	115.169	0.000		
Inverse normal Z	-1.043	0.149	-6.150	0.000		

Notes: Fisher-ADF test on 20 countries over years 1986-2012. H0: All panels contain unit roots. Ha: At least one panel is stationary. Country means are always removed. The test is performed with 3 lags on the levels, with 2 lags on the first differences and with 1 lag on the second-order difference.

Table A.11: Estimates with policies (EC specification)

Depvar: Δe	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Δy	0.508***	0.543***	0.525***	0.524***	0.460***	0.463***	0.440***
	(0.037)	(0.041)	(0.042)	(0.037)	(0.044)	(0.037)	(0.050)
Δy x $\tilde{grr}_{-1}$	0.563**			0.843***	0.991***	1.064***	1.127***
	(0.229)			(0.323)	(0.350)	(0.355)	(0.417)
Δy x $\tilde{train}_{-1}$		0.050**		0.030	0.062	0.032	0.031
		(0.020)		(0.026)	(0.041)	(0.025)	(0.039)
Δy x $\tilde{flex}_{-1}$			0.067**	0.113***	0.117***	0.066*	0.066
			(0.029)	(0.035)	(0.037)	(0.039)	(0.042)
Δy x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$					-0.123		-0.303
					(0.188)		(0.229)
Δy x $\tilde{grr}_{-1}$ x $\tilde{flex}_{-1}$					-0.829**		-0.625
					(0.331)		(0.416)
Δy x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					0.002		-0.018
					(0.041)		(0.046)
Δy x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					0.350		0.099
					(0.495)		(0.520)
e_{-1}	-0.104***	-0.106***	-0.112***	-0.119***	-0.130***	-0.152***	-0.127***
	(0.013)	(0.013)	(0.013)	(0.014)	(0.017)	(0.019)	(0.023)
e_{-1} x $\tilde{grr}_{-1}$	-0.013			-0.058	0.006	-0.005	0.034
	(0.035)			(0.041)	(0.066)	(0.058)	(0.072)
e_{-1} x $\tilde{train}_{-1}$		-0.001		0.000	-0.006	0.007**	-0.007
		(0.002)		(0.003)	(0.006)	(0.003)	(0.006)
e_{-1} x $\tilde{flex}_{-1}$			0.006	0.007	-0.003	-0.011	-0.017
			(0.005)	(0.006)	(0.008)	(0.008)	(0.010)
e_{-1} x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$					0.008		-0.066
					(0.036)		(0.046)
e_{-1} x $\tilde{grr}_{-1}$ x $\tilde{flex}_{-1}$					-0.006		-0.012
					(0.057)		(0.062)
e_{-1} x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					0.010*		0.008
					(0.005)		(0.006)
e_{-1} x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					0.048		0.017
					(0.045)		(0.052)
y_{-1}	0.071***	0.069***	0.071***	0.077***	0.091***	0.125***	0.123***
	(0.010)	(0.011)	(0.011)	(0.011)	(0.012)	(0.017)	(0.018)
y_{-1} x $\tilde{grr}_{-1}$	0.032			0.075**	0.024	0.008	-0.013
	(0.032)			(0.037)	(0.060)	(0.056)	(0.071)
y_{-1} x $\tilde{train}_{-1}$		-0.000		-0.002	0.001	-0.009**	0.002
		(0.002)		(0.003)	(0.005)	(0.004)	(0.006)
y_{-1} x $\tilde{flex}_{-1}$			-0.001	-0.002	0.005	0.020***	0.021**
			(0.004)	(0.005)	(0.007)	(0.007)	(0.009)

Table continues on the next page.

Depvar: Δe	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Continued from the previous page.							
y_{-1} x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$					-0.027		0.053
					(0.038)		(0.049)
y_{-1} x $\tilde{grr}_{-1}$ x $\tilde{flex}_{-1}$					-0.016		-0.003
					(0.052)		(0.057)
y_{-1} x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					-0.007		-0.004
					(0.005)		(0.005)
y_{-1} x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					-0.046		-0.008
					(0.042)		(0.049)
$\tilde{grr}_{-1}$	-0.433			-0.932**	-0.386	-0.125	0.080
	(0.359)			(0.423)	(0.673)	(0.639)	(0.812)
$\tilde{train}_{-1}$		0.004		0.023	-0.003	0.105***	-0.019
		(0.024)		(0.033)	(0.062)	(0.041)	(0.070)
$\tilde{flex}_{-1}$			0.003	0.005	-0.063	-0.242***	-0.248**
			(0.047)	(0.052)	(0.073)	(0.081)	(0.100)
$\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$					0.336		-0.580
					(0.429)		(0.557)
$\tilde{grr}_{-1}$ x $\tilde{flex}_{-1}$					0.217		0.085
					(0.580)		(0.642)
$\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					0.077		0.040
					(0.052)		(0.056)
$\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					0.507		0.081
					(0.468)		(0.539)
Additional LMIs						yes	yes
Additional LMIs x (Δy, e_{-1}, y_{-1})						yes	yes
Control variables	yes	yes	yes	yes	yes	yes	yes
Country dummies	yes	yes	yes	yes	yes	yes	yes
Year dummies	yes	yes	yes	yes	yes	yes	yes
Observations	502	502	502	502	502	502	502
R-squared	0.716	0.716	0.715	0.742	0.766	0.776	0.790
Wald test on joint significance of policy interactions:							
F-statistic					2.11		1.65
p-value					0.0073		0.0548

Notes: OLS estimation of (4) extended with interactions of e_{-1} and y_{-1} with the policy variables. Regressions in the last two columns also include additional LMIs and their interactions with Δy, e_{-1} and y_{-1}. Additional LMIs: uden, cove, cent, taxw, part. Control variables: first differences of trade openness, government size and working-age population growth. Robust standard errors are in parentheses. The policy variables are centered, i.e. expressed as deviations from their country means, and are lagged by one year. *** p<0.01, ** p<0.05, * p<0.1.

Table A.12: Estimates with policies (FD specification)

Depvar: Δe	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Δy	0.535***	0.546***	0.536***	0.539***	0.509***	0.492***	0.473***
	(0.041)	(0.044)	(0.044)	(0.042)	(0.049)	(0.043)	(0.054)
Δy x $\tilde{grr}_{-1}$	0.614**			0.842**	0.850**	1.104***	0.989**
	(0.247)			(0.381)	(0.394)	(0.393)	(0.432)
Δy x $\tilde{train}_{-1}$		0.061***		0.027	0.045	0.039	0.022
		(0.021)		(0.029)	(0.041)	(0.027)	(0.042)
Δy x $\tilde{flex}_{-1}$			0.047*	0.096***	0.077**	0.012	-0.025
			(0.027)	(0.034)	(0.036)	(0.037)	(0.041)
Δy x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$					-0.053		-0.309
					(0.195)		(0.240)
Δy x $\tilde{grr}_{-1}$ x $\tilde{flex}_{-1}$					-0.918***		-0.969**
					(0.351)		(0.415)
Δy x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					0.041		0.020
					(0.038)		(0.042)
Δy x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					0.154		0.027
					(0.497)		(0.528)
$\tilde{grr}_{-1}$	-0.015			-0.023	-0.030*	-0.028*	-0.032**
	(0.012)			(0.015)	(0.016)	(0.015)	(0.016)
$\tilde{train}_{-1}$		-0.001		-0.000	-0.002	-0.001	-0.001
		(0.001)		(0.001)	(0.001)	(0.001)	(0.001)
$\tilde{flex}_{-1}$			0.002	0.000	0.002	0.001	0.003
			(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
$\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$					0.001		0.009
					(0.009)		(0.009)
$\tilde{grr}_{-1}$ x $\tilde{flex}_{-1}$					-0.001		0.014
					(0.014)		(0.015)
$\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					-0.002		-0.002
					(0.002)		(0.002)
$\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$					-0.023		-0.023
					(0.017)		(0.018)
Additional LMIs						yes	yes
Additional LMIs x Δy						yes	yes
Control variables	yes	yes	yes	yes	yes	yes	yes
Country dummies	yes	yes	yes	yes	yes	yes	yes
Year dummies	yes	yes	yes	yes	yes	yes	yes
Observations	502	502	502	502	502	502	502
R-squared	0.669	0.672	0.667	0.684	0.696	0.703	0.715
Wald test on joint significance of policy interactions:							
F-statistic					2.22		1.87
p-value					0.0249		0.0623

Notes: OLS estimation of (4), including additional LMIs and their interactions with Δy in the last two columns. Additional LMIs: uden, cove, cent, taxw, part. Control variables: first differences of trade openness, government size and working-age population growth. Robust standard errors are in parentheses. The policy variables are centered, i.e. expressed as deviations from their country means, and are lagged by one year. *** p<0.01, ** p<0.05, * p<0.1.

Table A.13: Effects on the short-run elasticity of σ-increases in the policy variables – FD specification

	baseline				with additional LMIs			
Policy variable	without interactions	with interactions			without interactions	with interactions		
		average	liberal	interventionist		average	liberal	interventionist
grr	0.103**	0.104**	0.018	0.165**	0.135***	0.121**	0.051	0.170**
	(0.047)	(0.048)	(0.056)	(0.069)	(0.048)	(0.053)	(0.073)	(0.070)
train	0.034	0.056	0.057	0.006	0.049	0.028	0.219	-0.020
	(0.036)	(0.051)	(0.174)	(0.048)	(0.034)	(0.053)	(0.208)	(0.049)
flex	0.091***	0.073**	0.130***	0.046	0.011	-0.024	0.042	-0.077
	(0.032)	(0.034)	(0.048)	(0.058)	(0.035)	(0.039)	(0.058)	(0.065)

Notes: The figures show the effects of a 1-standard-deviation increase in a policy variable on the short-run employment-output elasticity. Baseline estimates are based on columns 4 (without interactions) and 5 (with interactions), those with additional LMIs on column 6 and 7 of Table A.12. Average, liberal and interventionist regimes differ in the values the non-changing policy variables are assumed to take. Standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A.14: Effects on the short-run elasticity of σ-increases in the policy variables – Spain included

	baseline				with LMIs incl. temp			
Policy variable	without interactions	with interactions			without interactions	with interactions		
		average	liberal	interventionist		average	liberal	interventionist
grr	0.123***	0.109**	0.056	0.123*	0.155***	0.170***	0.195*	0.183**
	(0.039)	(0.045)	(0.059)	(0.068)	(0.051)	(0.064)	(0.107)	(0.075)
train	-0.023	-0.023	0.159	-0.126**	0.019	-0.024	0.364	0.031
	(0.036)	(0.058)	(0.173)	(0.060)	(0.042)	(0.049)	(0.295)	(0.064)
flex	0.050	0.027	0.055	0.042	0.181***	0.171***	0.200**	0.075
	(0.035)	(0.039)	(0.052)	(0.057)	(0.056)	(0.061)	(0.100)	(0.074)

Notes: The figures show the effects of a 1-standard-deviation increase in a policy variable on the short-run employment-output elasticity. The baseline regression is as in Table 4, except that the sample includes Spain. The regression with additional LMIs also includes the share of temporary employees (temp). The inclusion of temp reduces the sample size from N=529 (baseline) to N=365. Average, liberal and interventionist regimes differ in the values the non-changing policy variables are assumed to take. Standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A.15: Effects on the long-run elasticity of σ-increases in the policy variables

Policy variable	without interactions	with interactions		
		average	liberal	interventionist
grr	0.037***	0.027	0.025	0.028*
	(0.012)	(0.017)	(0.019)	(0.016)
train	-0.017	-0.026	-0.015	-0.027
	(0.012)	(0.019)	(0.039)	(0.025)
flex	0.026*	0.019	0.027	0.000
	(0.014)	(0.014)	(0.026)	(0.017)

Notes: The figures show the effects of a 1-standard-deviation increase in a policy variable on the long-run employment-output elasticity. Estimates without interactions are based on column 4, those with interactions on column 5 of Table A.11. Average, liberal and interventionist regimes differ in the values the non-changing policy variables are assumed to take. Standard errors (in parentheses) are calculated with the delta method. *** p<0.01, ** p<0.05, * p<0.1.

Table A.16: Policy distance from Denmark in 2010

Country code (ISO)	Policy variable grr	train	flex
AUS	-0.210	-2.973	0.481
AUT	-0.097	0.149	-0.111
BEL	0.006	-1.455	-0.387
CAN	-0.273	-1.838	1.177
CHE	-0.084	-0.606	0.419
DEU	-0.182	-0.894	-0.161
DNK	0.000	0.000	0.000
ESP	-0.072	-2.494	-0.807
FIN	-0.081	-0.461	-0.097
FRA	-0.019	-0.896	-1.083
GBR	-0.251	-3.892	0.926
IRL	0.014	-0.989	0.817
ITA	-0.069	-1.623	-0.626
JPN	-0.279	-2.860	0.655
NLD	-0.076	-1.129	-0.218
NOR	-0.078	-0.120	-0.793
NZL	-0.172	-2.037	0.492
PRT	0.024	-1.383	-1.399
SWE	-0.026	-2.133	-0.041
USA	-0.146	-2.865	1.564

Notes: Absolute deviations of the policy variables from the Danish levels in year 2010.

Table A.17: Impact of flexicurity reform: EC specification

	(1) EC baseline			(2) EC additional LMIs			(3) EC broad ALMP		
	sre change	s.e.		sre change	s.e.		sre change	s.e.	
AUS	0.283	0.096	***	0.263	0.115	**	0.252	0.083	***
JPN	0.261	0.072	***	0.304	0.122	**	0.253	0.083	***
DEU	0.244	0.073	***	0.198	0.081	**	0.246	0.070	***
GBR	0.207	0.090	**	0.267	0.131	**	0.173	0.089	*
ITA	0.201	0.067	***	0.113	0.077		0.198	0.077	***
NZL	0.200	0.081	**	0.174	0.094	*	0.233	0.097	**
NOR	0.160	0.066	**	0.082	0.071		0.149	0.069	**
NLD	0.149	0.053	***	0.091	0.059		0.090	0.036	**
FRA	0.129	0.073	*	0.026	0.080		0.156	0.082	*
SWE	0.128	0.089		0.037	0.093		0.060	0.044	
FIN	0.104	0.037	***	0.073	0.041	*	0.117	0.045	***
AUT	0.085	0.044	*	0.069	0.043		0.110	0.041	***
BEL	0.069	0.061		-0.009	0.066		0.066	0.066	
CAN	0.063	0.071		0.137	0.132		0.105	0.064	
CHE	0.061	0.038		0.060	0.042		0.053	0.039	
PRT	0.057	0.088		-0.049	0.096		0.041	0.090	
USA	0.054	0.142		0.045	0.150		0.051	0.159	
IRL	-0.001	0.066		-0.005	0.067		-0.013	0.102	

Notes: Estimated changes in the short-run employment-output elasticity and the corresponding standard errors from three different EC regressions. The calculations behind the estimates in the first two columns are based on columns 5 and 7 of Table A.11. The figures in the last column are derived from an EC regression where train is replaced by a broad ALMP variable. Countries are listed from the highest to the lowest change according to the baseline estimates. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.18: Impact of flexicurity reform: FD specification

	(1) FD			(2) FD country trends			(3) FD per-capita transf		
	sre change	s.e.		sre change	s.e.		sre change	s.e.	
AUS	0.252	0.097	***	0.261	0.098	***	0.252	0.092	***
JPN	0.241	0.076	***	0.275	0.076	***	0.233	0.074	***
DEU	0.219	0.082	***	0.197	0.080	**	0.172	0.079	**
GBR	0.253	0.090	***	0.321	0.096	***	0.255	0.092	***
ITA	0.143	0.068	**	0.187	0.071	***	0.107	0.065	*
NZL	0.177	0.084	**	0.179	0.084	**	0.165	0.079	**
NOR	0.160	0.074	**	0.128	0.073	*	0.135	0.068	**
NLD	0.122	0.056	**	0.134	0.057	**	0.096	0.053	*
FRA	0.079	0.077		0.119	0.079		0.063	0.070	
SWE	0.097	0.093		0.144	0.094		0.076	0.089	
FIN	0.095	0.042	**	0.091	0.041	**	0.070	0.040	*
AUT	0.090	0.050	*	0.071	0.047		0.058	0.049	
BEL	0.034	0.065		0.098	0.066		0.014	0.063	
CAN	0.058	0.082		0.119	0.078		0.045	0.079	
CHE	0.054	0.040		0.058	0.041		0.033	0.039	
PRT	-0.026	0.088		0.116	0.094		-0.060	0.088	
USA	0.144	0.139		0.128	0.149		0.172	0.129	
IRL	0.025	0.069		0.020	0.071		0.021	0.068	

Notes: Estimated changes in the short-run employment-output elasticity and the corresponding standard errors from three different FD regressions. The figures in the first column are based on the estimations in column 5 of Table A.12. Countries are listed from the highest to the lowest change according to the baseline estimates from the EC regression presented in the first column of Table A.17. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A.19: Two-step estimation with FMOLS and DOLS

1st stage: Cointegrating relationship				
	FMOLS	FMOLS	DOLS	DOLS
	cons	cons + trend	cons	cons+trend
y	0.297***	0.144***	0.311***	0.186***
	(0.011)	(0.031)	(0.012)	(0.030)
Observations	1040	1040	1033	1032
2nd stage: Dynamic relationship				
Δy	0.323***	0.307***	0.320***	0.314***
	(0.028)	(0.029)	(0.029)	(0.030)
ECM_{-1}	-0.034***	-0.037***	-0.028***	-0.044***
	(0.004)	(0.012)	(0.005)	(0.013)
Observations	1,020	1,020	1,013	1,012
R-squared	0.509	0.487	0.498	0.492

Notes: First-stage dependent variable is the log number of employees, second-stage dependent variable is the log annual change in the number of employees. The sample consists of 20 countries over years 1960-2012. First-stage estimations are performed in EViews using default options for the assumptions concerning the long-run variance. DOLS includes zero or one lead and lag, depending on the Akaike IC. The second-stage regression is estimated with country fixed effects and year dummies. Standard errors are reported in parentheses. *** p<0.01, ** p<0.05, * p<0.1.