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Holger Görg
Kiel Centre for Globalization, Kiel Institute for the World Economy, University of Kiel and IZA

Cecília Hornok
Kiel Centre for Globalization and Kiel Institute for the World Economy

Catia Montagna
University of Aberdeen and the Scottish Institute for Research in Economics

George E. Onwordi
University of Aberdeen

NOVEMBER 2018

IZA – Institute of Labor Economics

Schaumburg-Lippe-Straße 5–9
53113 Bonn, Germany

Phone: +49-228-3894-0
Email: publications@iza.org

www.iza.org

ABSTRACT

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

How do labour market policies influence employment's responsiveness to output fluctuations (employment-output elasticity)? We revisit this question on a panel of OECD countries, which also incorporates the period of the Great Recession. We distinguish between passive and active labour market policies and allow for their interactions, i.e. the policy mix, to play a role. We find that the effects of any single policy change are shaped by the broader existing policy-mix within which it takes place. Finally, we evaluate the effect of a move to 'flexicurity' on the employment-output elasticity in each country.

JEL Classification: E24, E32, J21, J65

Keywords: employment-output elasticity, labour market policy, welfare state, flexicurity

Corresponding author:
Holger Görg
Kiel Institute for the World Economy
Kiellinie 66
24105 Kiel
Germany
E-mail: holger.goerg@ifw-kiel.de

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1 Introduction

The response of employment to external shocks is a central concern to policy-makers. Crivelli *et al.* (2012), e.g., 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 this inter-country variation in the employment responsiveness to output shocks (Crivelli *et al.*, 2012; Bassanini, 2012; Blanchard and Wolfers, 2000).[1]

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 relationship. Secondly, and most importantly, we shall 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. This is important because the interaction of these two policy dimensions underpins an important debate about how labour market reforms in the direction of the *flexicurity* system, for instance as practised in Denmark, can affect aggregate labour market outcomes.

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 system with flexible hiring and firing rules and high levels of ALMPs 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. 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 qualititative assessments (e.g. Andersen and Svarer, 2007) or on the effect of specific aspects of the flexicurity package (e.g. Card *et al.*, 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 *et al.*

[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 *et al.*, 2018b), 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 *et al.*, 2018a).

(2016) and Molana *et al.* (2018b).

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 2014 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 has increased over recent decades, while the long-run elasticity remained stable.

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 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 – *average*, *flexible* and *generous* – of labour market regimes. The former is a hypothetical system that we obtain by setting the policy variables to their sample average values. The *flexible* regime, such as that of 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), set at their 25th percentiles. The *generous* regime, in contrast, is characterized 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 *et al.* (2018a, 2018b), 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 flexible regime is not associated with higher short-run employment-output elasticity. In a generous 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 very flexible regimes, the effect of flexicurity reform on employment's responsiveness is not statistically significant. However, we find that labour market reforms toward flexicurity would imply greater short-run employment-output elasticity in, e.g., Great Britain and Germany in the order of approximately 0.25 percentage points. Interestingly, while Germany is both less generous and less flexible compared to Denmark (our benchmark) and have high levels of expenditure on training programmes, Great Britain is much less generous and more flexible, and has 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, and may well lead to 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 *et al.*, 2002; Blanchard and Wolfers, 2000; Crivelli *et al.*, 2012; Faccini and Bondibene, 2012; Nickell *et al.*, 2005). Some of these works only look at the long-run impact of these policies on the labour market – with the exception of Bassanini (2012), Crivelli *et al.* (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 complementaries 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 *et al.* (2015) examine the effects of labour market institutions in driving business cycle fluctuations. Fonseca *et al.* (2010) also explore the relationship between labour market institutions and business cycle, focusing on international co-movements. The effects of institutions on the dynamic adjustment of inflation has 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 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 a 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 quantitative assessment of the effectiveness of ALMP measures, concluding that training programmes have had mild effects on employment outcomes. Shahidi *et al.* (2016) assesses 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, while Section 4 obtains the baseline elasticity estimates. 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. Finally, Section 7 evaluates the flexicurity reform and Section 8 concludes.

2 Empirical approach

Our starting point is a general empirical model that describes the dynamic relationship between employment and output:[2]

$$\Delta e_{i,t} = \alpha e_{i,t-1} + \beta y_{i,t-1} + \gamma \Delta y_{i,t} + \varepsilon_{i,t}, \tag{1}$$

where $\Delta e_{i,t}$ denotes the (annual) change of employment in country i at time t which is driven by the annual change of output ($\Delta y_{i,t}$) and the lagged values of these two variables in levels. All variables are expressed in natural logarithms. $\varepsilon_{i,t}$ is the error term that may include time-invariant country as well as aggregate time effects.

For ease of interpretation, equation (1) can be expressed as

$$\Delta e_{i,t} = \alpha \left[e_{i,t-1} + \frac{\beta}{\alpha} y_{i,t-1} + \text{trend} \right] + \gamma \Delta y_{i,t} + \varepsilon_{i,t},$$

where the term in brackets captures the long-run cointegrating relationship between employment and output (given that it exists) and also allows for a deterministic trend. The above implies that any deviation in employment from its long-run equilibrium will result in an error-correction adjustment process, where the speed of adjustment is given by α ($\alpha < 0$).[3] This empirical approach, which does not rely on estimating aggregate labour demand or labour supply equations separately,

[2] A similar empirical approach is taken for the estimation of the employment-output elasticity by Behar (2015).

[3] Gnocchi *et al.* (2015) employ a similar approach to examine the direct relationship between labour market policy instruments and the business cycle.

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.

The short-run employment-output elasticity is captured by γ, while its long-run counterpart is $-\frac{\beta}{\alpha}$. Both are positive, since β and γ are positive and α is expected to be negative. Should the estimated value of α be non-negative, then the cointegrating relationship would not exist (Ericsson and MacKinnon, 2002). In this case, equation (1) would simplify to

$$\Delta e_{i,t} = \gamma \Delta y_{i,t} + \varepsilon_{i,t}, \tag{2}$$

and the short- and long-run elasticities would both be γ. In what follows, we refer to equation (1) as the error correction specification and equation (2) the first-difference spefication.

To incorporate the role of policies we follow on previous literature (e.g. Blanchard and Wolfers, 2000; Crivelli *et al.*, 2012; Bassanini, 2012) and include interactions of the right-hand side variables with the policy variables in the estimating equation. The first-difference specification will then become

$$\Delta e_{i,t} = \gamma_1 \Delta y_{i,t} + \gamma_2 \Delta y_{i,t} \tilde{\mathbf{P}}_{\mathbf{i,t-1}} + \eta \tilde{\mathbf{P}}_{\mathbf{i,t-1}} + \varepsilon_{\mathbf{i,t}}, \tag{3}$$

where $\tilde{\mathbf{P}}_{\mathbf{i,t-1}}$ denotes the vector of policy variables (and possibly their interactions with each other). To overcome the concern of policy variables being endogeneous to the contemporaneous employment dynamics, we capture policies with their lagged values. Also, for the ease of interpretation, the policy variables are 'centered' within each year, i.e. expressed as deviations from their country means as $\tilde{P}_{i,t} = P_{i,t} - \frac{1}{N}\sum_{j=1}^{N} P_{j,t}$.

In the above formulation the short-run elasticity of any given country can be interpreted as an average elasticity plus country deviations originating from policy differences at any given point in time,

$$\text{sre}_i = \gamma_1 + \sum_k \gamma_2^k \tilde{P}_i^k \tag{4}$$

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.

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

The error correction specification (1) can similarly be extended to incorporate the policy variables. Specifically, doing so will result in estimating an equation such as (3) but augmented with the terms $\alpha_1 e_{i,t-1} + \alpha_2 e_{i,t-1}\mathbf{\tilde{P}_{i,t-1}}$ and $\beta_1 y_{i,t-1} + \beta_2 y_{i,t-1}\mathbf{\tilde{P}_{i,t-1}}$. 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 elasticity, the long-run elasticity can be expressed as a function of parameters common to all countries and policy-induced individual deviations, more formally, $-\frac{\beta_1+\sum_k \beta_2^k \tilde{P}_i^k}{\alpha_1+\sum_k \alpha_2^k \tilde{P}_i^k}$.

3 Data and time series properties

Our database is a panel of OECD countries with the longest possible time dimension. We start with a balanced panel of employment and output for 20 countries and 55 years (1960-2014). 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 Spain), Australia, Canada, Japan, New Zealand, Norway, Switzerland, and the US. We measure employment and output with, respectively, total employment in persons and the volume of GDP – both of which are sourced from the Penn World Table.[5] A detailed description of the policy variables is provided in Section 5, while further definitions and data sources are summarized in Table A.1 in the Appendix.

Before we delve into the econometric analysis, we consider the time series properties of our data. In particular, we carry out unit root tests on the balanced panel of 55 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-county tests and in Tables A.4 and A.5 for the panel. Based on these test results we conclude that, in levels, both employment and output have unit roots against being trend stationary, while their first differences can be considered stationary for the majority (though not all) of the countries.

More specifically, the test results are the following. The country ADF tests cannot reject the null hypothesis of a unit root for the levels (except for output in Switzerland), while they always reject for the first differences. The KPSS test rejects the null of stationarity for the levels in most of the cases (also including output in Switzerland). The results for first differences are somewhat more

[5]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`.

mixed, rejecting stationarity in a number of 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, which always rejects the null of unit root for the first differenes.

The panel unit root tests reinforce 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 though, with null hypothesis of all panels being stationary and alternative hypothesis of at least some of the panels containing a unit root, rejects the null, indicating that for some countries the first differences cannot be considered stationary.

With these results in mind, we continue with the assumption that output and employment have unit roots, while their first differences are stationary.

This does not tell us, however, whether the cointegrating relationship in model (1) exists, which we test in two ways. First, we perform the Johansen cointegrating rank test on each country (Johansen, 1991), also 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 one cointegrating equation between employment and output. Where no cointegration is found, the test statistics are often not far from the critical values. When the number of cointegrating equations is chosen by minimizing the HQIC (Hannan and Quinn Information Criterion), we find cointegration for all countries but Ireland.

Second, in the next section, we test for the presence of cointegration by testing the statistical significance of the speed of adjustment parameter when estimating equation (1) on the pooled 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 show, the t-statistics for the speed of adjustment parameter estimate (in brackets) lie outside the 5% critical values, indicating the existence of cointegration.

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 error correction and the first-difference models and indicate whenever the presumption of cointegration leads to qualitatively different results.

4 Baseline elasticity estimates

How responsive was employment to output fluctuations in the past decades in developed OECD countries? Assuming a permanent output change, did an immediate response differ substantially from the long-run effect? To answer these questions, we estimate models (1) and (2) on our panel of 20 countries, using the full time span between 1960 and 2014. Since, due to data availability, our later inclusion of policy variables in the analysis requires a shorter time dimension, we also consider the period 1986-2012. All regressions include a full set of country and year dummies and are estimated with Ordinary Least Squares and robust standard errors. Table 1 reports the results.

Table 1: Elasticity estimates

Depvar: Δe	1960-2014		1986-2012	
Δy (Short-run elasticity)	0.305***	0.333***	0.545***	0.546***
	(0.031)	(0.030)	(0.046)	(0.045)
e_{-1} (Speed of adjustment)		-0.034***		-0.062***
		(0.004)		(0.014)
		[-9.11]		[-4.53]
y_{-1}		0.028***		0.048***
		(0.005)		(0.013)
Country dummies	✓	✓	✓	✓
Year dummies	✓	✓	✓	✓
Observations	1,080	1,080	540	540
R-squared	0.469	0.510	0.636	0.653
Long-run elasticity		0.819***		0.778***
		(0.138)		(0.149)
CI test 5% critical value		-3.96		-4.15

Notes: Estimation of (1) and (2) with OLS. Parentheses include robust standard errors, brackets t-statitics. 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.

The short-run employment-output elasticity of the average country in our sample is estimated to be 0.3 on the long panel and 0.5 on the shorter one. The estimate of 0.5 means 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 indicates that, for the average country, employment has become more responsive to output shocks over the recent decades. The error correction specification allows us to estimate the long-run employment-output elasticity, which is around 0.8 for the average country on both samples, meaning that if output permanently increases by 1%, employment will be up by 0.8% in the long run. Figure A.1 in the Appendix illustrates the dynamic adjustment of employment to a 1% permanent output shock over time. As is evident from the figure, the instantanous impact is greater in the shorter panel than in the longer one. However, the adjustment of employment as the impact of the increase in output dies out is the same for both panels, essentially resulting in similar long-run elasticities.[6]

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

To investigate more closely how the short-run elasticity changes over time, we estimate equation (1) on the long panel by interacting the output change variable with (five-year) period dummies and plot the period-specific elasticity estimates on Figure 1.[7] The short-run elasticity has an upward trend, which has clearly become steeper from the second half of the 1980s. While in the first two and a half decades it was, on average, close to 0.2, it rose quickly to around 0.5 afterwards. The reasons for this increase are likely structural and, as Blanchard *et al.* (2006) suggest, might also 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.

In order to establish the extent of inter-country heterogeneity in employment responsiveness, we estimate model (1) by interacting the output change variable with country dummies to obtain country-specific elasticities. Figure 2 reports the short-run elasticities by country, estimated either on the long panel or the short panel, and plotted in descending order of the elasticity obtained from the long panel. It is apparent that there is considerable inter-country variation in the responsiveness of employment to output. Countries with the largest elasticities at the top of the graph (Ireland, US) have at least three times higher employment responsiveness than countries at the bottom (e.g.

[6] A response function on Figure A.1 is 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.

[7] Estimating the first-difference specification (2) leads to very similar results.

Figure 1: Short-run elasticity estimates over time

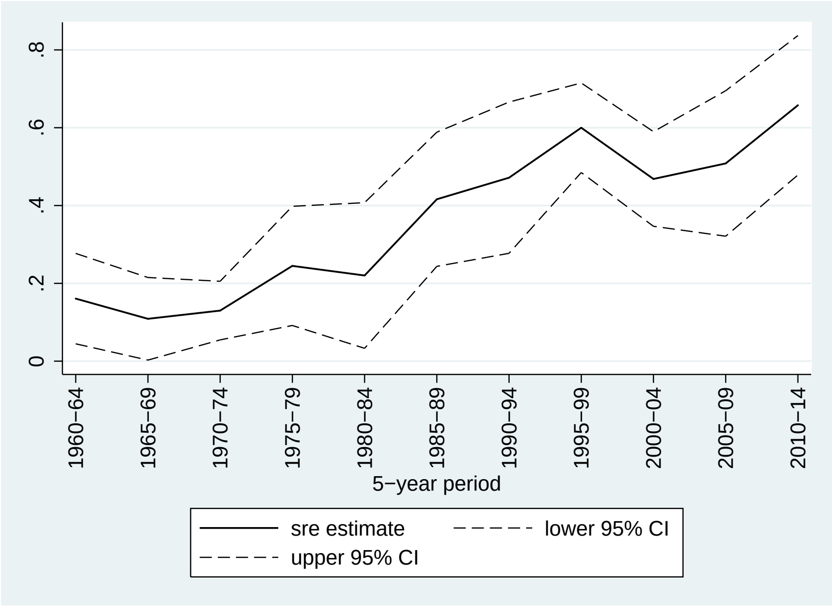
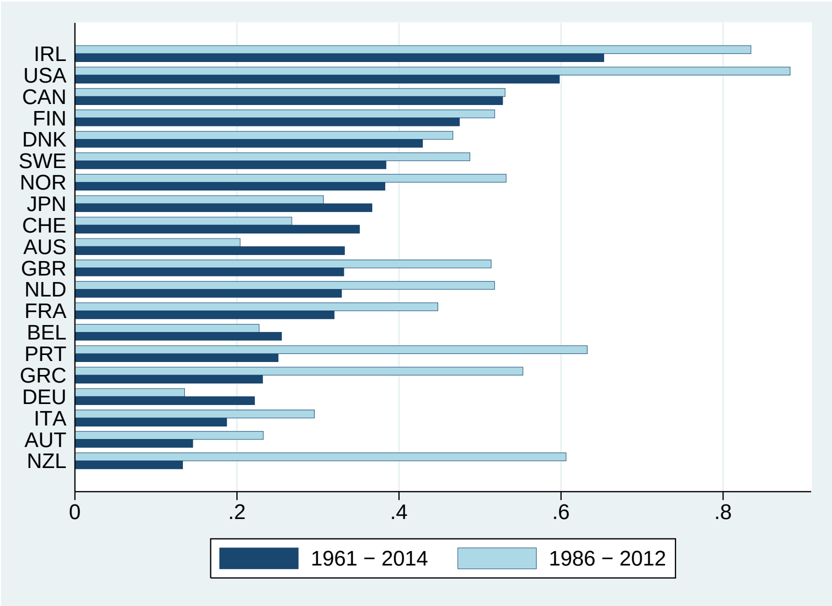


Figure 2: Estimated short-run elasticities by country



Austria, Germany).[8] The estimates for some countries also change considerably when the sample is shortened to the more recent decades. In particular, the elasticities of Greece, New Zealand and Portugal, which all have below-average elasticities in the long panel, increase to above average.

[8]We obtain qualitatively the same results by estimating the first-difference specification, although we get a somewhat different country ranking.

These country differences (and possibly also their changes over time) might reflect – among others things – differences in labour market institutions and policies, which we investigate in the remaining sections of the paper.

5 Labour market policies in OECD countries

We characterise the labour market regime of a country with three policy measures: (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 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 pillar of *flexicurity* – which has been widely endorsed as a means to reconciling the needs for the flexibility required to adjust to negative shocks with that for adequate social safety nets.

Table 2 reports descriptive statistics of the three policy variables.

Table 2: Descriptives of policy variables

Variable	Description	N	Mean	Std.Dev.	Min	Max
grr	Gross unemployment replacement rate	516	0.304	0.125	0.025	0.652
flex	Labour market flexibility index (0 - 6)	516	4.027	0.987	1.677	5.746
train	Training expenditure per unemployed (log)	516	7.151	1.125	3.150	9.252

Notes: The sample covers the period 1985-2011, for which all the three variables are available. Training expenditures are adjusted for purchasing power parity differences and expressed in the natural logarithm of PPP US dollars. All three variables are calculated based exclusively on OECD data.

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.[9] *grr* is a summary measure, calculated by the OECD as an average for two earnings levels, three family situations and three durations of unemployment.[10] 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.

[9] 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.

[10] 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 using the OECD Tax-Benefit Models. To minimize the impact of this methodological break, we rely on the figures until 2005 and extend the time series using the growth rates of the following years.

The benefit replacement rate of the average country in our sample is approximately 0.3, i.e. the unemployment benefit is close to one-third of the previous gross earnings. 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.2). 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.

Next, we capture labour market flexibility (*flex*) with the inverse of the Employment Protection Index (EPL) of the OECD, available for the years 1985-2013.[11] Our *flex* index measures, on a 0 to 6 scale, how easy it is for employers to lay off workers or to employ them on temporary work contracts, as it is laid down by a country's regulatory system.[12] Hence, a higher value of *flex* means a more flexible labour market. Table 2 shows that the value of the flexibility index is 4 on average in our sample and varies between 1.7 and 5.7. The most flexible labour markets are in the US, Canada and Great Britain, while the Portuguese and the Greek markets are the least flexible. Time series graphs on Figure A.3 reveal a tendency among countries with less flexible labour markets to increase flexibility over time, which especially applies to Germany, Sweden, Italy, Portugal and, more recently, Greece.

Finally, 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 *targetted* on unemployed persons or employees at high risk of becoming unemployed. Training programmes which are generally available are considered to be part of the general education and therefore excluded from this definition. Although the full spectrum of ALMPs include 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 promoting workers' employability. Second, training makes up 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

[11] A description of the EPL index is provided in OECD (2013).

[12] Following the OECD practice (OECD, 2013), we take the weighted average of the employment protection sub-indices for the dismissals from regular contracts (EPRC) and the use of temporary contracts (EPT) in the following way: $\text{EPL} = \frac{7}{12}\text{EPRC} + \frac{5}{12}\text{EPT}$. Then we take the inverse (6-EPL) to create our flex varible.

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 7.151, i.e. 1,275 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 relative high importance of ALMPs within Europe (Figure A.4).

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.7 reveal that countries with generous unemployment benefits also spend more on training schemes. This positive relationship is highly statistically significant and has been getting 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 (Nickell *et al*, 2005). Furthermore, generous systems also tend to coexist with less flexible markets, though this relationship becomes statistically significant only in the more recent years of our sample period.

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

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. Simply cross-plotting the sample means of the policy variables with the short-run country elasticities (estimated on the short panel) from Section 4 reveals no clear correlation patterns (Figure A.5). Our regression results will however show that at least two of the three policy measures – benefit generosity and labour market flexibility – are important determinants of the employment-to-output responsiveness.

We estimate equation (3) and its error-correction counterpart and report the results in Table A.8 and Table A.9, respectively. We introduce the policy variables and their interactions with the right-hand-side variables sequentially and find that the coefficient estimates are quite stable across the different specifications. In the last two columns of the tables we also control for other possible covariates of employment growth, such as the annual change of openness to international trade, union density and the size of the government (measured as the share of government consumption in GDP).[13] The inclusion of these control variables - although it improves the model fit significantly - leaves the main results unchanged.

The first rows in the tables show the short-run employment-output elasticity estimate for the average country (around 0.5), while the coefficients of the interaction terms capture how the elasticity varies due to deviations from the average policy. All things being equal, a country with an unemployment replacement rate that is higher than the average by 0.1 (i.e. 10 percentage points) tends to have a 0.0751 higher-than-average short-run elasticity, as shown by the coefficient of 0.751 in column 6 of Table A.9. Similarly, a one-point higher flexibility index corresponds to a 0.089 higher elasticity. In contrast, training expenditure is not found to have a significant explanatory power for the short-run elasticity.

Next, we consider all possible interactions between the policy variables (columns 5 and 7 of Table A.8 and Table A.9). Broadly, our results show that policies do not only matter individually but also in combination. For instance, we find that although when taken individually, higher benefit generosity and more labour market flexibility both associate with a higher employment responsiveness, there is an offsetting effect when the two are considered jointly. Also, the interaction of *train* and *flex* turns out to have a positive influence on employment elasticity, which is though statistically significant only in the first-difference estimation. Testing the joint significance of the terms with policy interactions also confirms that accounting for policy complementarities significantly improves the model fit.[14] 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.

By and large, the results reported above are consistent with theoretical predictions. From the perspective of search and matching framework, a higher generosity of unemployment benefit increases the responsiveness of employment to shocks (Molana *et al.*, 2018a; Shimer, 2005; Zanetti, 2011) as our estimates reveals. The driving force is that a higher benefit raises the outside option of workers and their threat point in wage bargain, with attendant reduction in match profitability. This in turn makes firms more 'sensitive' and implies that a shock that affects productivity will

[13] The control variables are described in Table A.1. The annual changes of these variables are stationary time series. Unit root test results are avaible on request.

[14] The test results are available on request.

result in greater movement (in percentage terms) in job 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 flexibililty 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 work towards higher employment responsiveness; on the other hand, increased flexibility in terms of firing rules can have a dampening effect on employment fluctuations. Thus, when both policies are implemented jointly (i.e. via increases in both hiring and firing flexibility), the impact on employment fluctuation 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). 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 (e.g. Onwordi, 2016; Pissarides, 2009; Stähler and Thomas, 2012), a positive shock to productivity is likely to result in greater elacticity in the short-run. However, as argued by Molana *et al.* (2018a), 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 exogenous shocks.

To further investigate the impact of these policies, we study their complementarities more systematically. Particularly, we examine the effects of, ceteris paribus, 1-standard-deviation upward moves in each policy variable under three labour market regimes: *average*, *flexible* and *generous*. 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 (non-centered averages are reported in Table 2). The *flexible* regime assumes a flexibility index at the 75th percentile of its sample distribution ($flex$ = 4.8, non-centered) and the other two policy variables at their 25th percentiles (grr = 0.2, $train$ = 6.6, both non-centered). The *generous* regime, in contrast, is characterized by a flexibility index at its 25th percentile ($flex$ = 3.3) and the other two at their 75th percentiles (grr = 0.4, $train$ = 7.9). These inter-quartile differences roughly coincide with the average differences between the typical flexible and generous countries in our sample.

We rely on expression (4) to calculate the effects on the short-run elasticity, using the estimated coefficients in Table A.9, reported either in column 6 (without policy interactions) or in column

7 (with policy interactions).[15] Without policy interactions, the effect on the short-run elasticity of a 1-standard-deviation upward move in the gross replacement rate, for instance, is obtained as $\hat{\gamma}_2^{\tilde{\text{grr}}} \times \hat{\sigma}^{\text{grr}} = 0.751 \times 0.125 = 0.094$, where $\hat{\gamma}_2^{\tilde{\text{grr}}}$ is the estimated coefficient for the interaction of Δy with the gross replacement rate in column 6 of Table A.9, and $\hat{\sigma}^{\text{grr}}$ is the sample standard deviation of the gross replacement rate (Table 2). 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 7 of Table A.9 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 in Table 3 together with standard errors which reflect parameter uncertainty.

Table 3: Effects on the short-run elasticity of σ-increases in the policy variables

Policy variable	without interactions	with interactions		
		average	flexible	generous
grr	0.094***	0.081**	0.005	0.197***
	(0.034)	(0.036)	(0.055)	(0.059)
train	0.000	0.041	0.059	0.062
	(0.033)	(0.033)	(0.141)	(0.044)
flex	0.088***	0.090**	0.125***	0.034
	(0.029)	(0.037)	(0.037)	(0.054)

Notes: The figures show the effects of a 1-standard-deviation increase in a policy variable on the short-run employment-output elasticity. Estimates without interactions are based on column (6), those with interactions on column (7) of Table A.9. *Average*, *flexible* and *generous* regimes differ in the values the non-changing policy variables are assumed to take. Robust 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 of Table 3): a move towards more unemployment benefit generosity and a more flexible labour market both correspond to larger short-run employment-output elasticity (by 0.094 and 0.088, respectively). When, however, policy complementarities are taken into account, it matters considerably in which labour market regime a certain policy change takes effect. In *flexible* regimes, making the unemployment benefit more generous does not result in significantly higher short-run elasticity. In a similar vein, a reform towards more labour market flexibility does not

[15]The results based on the coefficients in Table A.8 are qualitatively the same and available on request.

lead to a significantly higher short-run employment responsiveness under a *generous* policy mix. Meanwhile, the estimated effects under the *average* regime are, as expected, very close to those in the first column, where complementarities are not taken into account. An important implication of these results is, therefore, that the impact of a specific policy on employment fluctuations depends on the initial policy mix.

Finally, we also calculate the corresponding changes in the long-run elasticity, which we report in Table A.10 in the Appendix.[16] The numbers suggest that the effects of these policies found on the short-run elasticity mostly fade out in the long run. Training is the only policy with a (weakly) significant and negative long-run impact on the employment elaticities in more generous regimes. This 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 *et al.*, 2018a).[17]

7 Reforming toward 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 characterized 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-know 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.11 in the Appendix displays the countries' policy distances from Denmark in 2010. Clearly, these figures imply that a reform in the direction of flexicurity would require in most of the countries within our sample to increase their benefit generosity and raise their expenditures on training, while, at the same time, adjust the flexibility of their labour markets toward an average level. In light of this, our priors are that a reform in the direction of flexicurity is likely to increase

[16]This we do by evaluating the expression $-\frac{\hat{\beta}_1+\sum_k \hat{\beta}_2^k \tilde{P}^k}{\hat{\alpha}_1+\sum_k \hat{\alpha}_2^k \tilde{P}^k}$ before and after the one-stdev increase in a given policy variable and taking the difference. For the calculation 'without interactions' we assume that policy variables take their average values. Taking other values does not influence the results qualitatively.

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

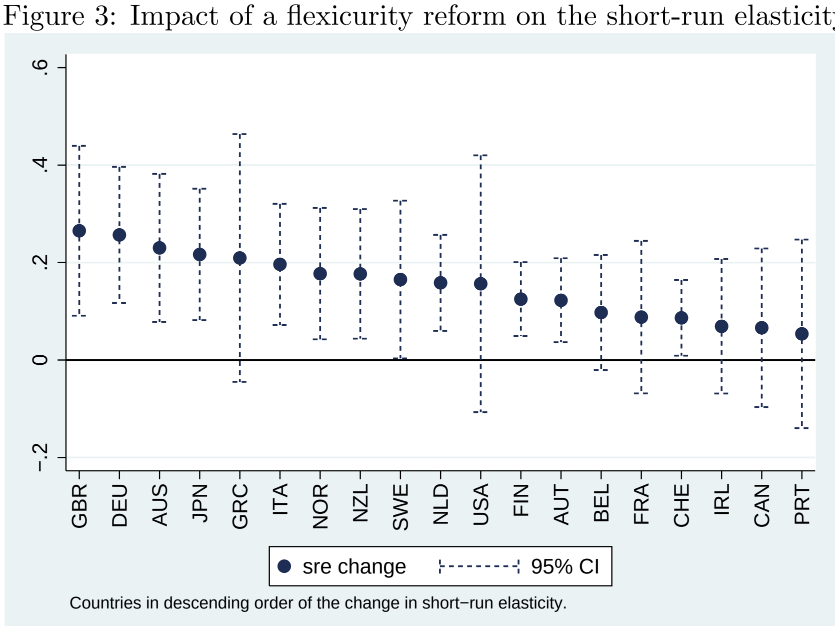
or leave the short-run elasticity of all countries roughly unchanged. The main reason, as suggested by our data, is that this reform would make each country's regime more generous or leave them almost as generous as they were. Moreover, for most countries with regimes characterised by a generosity similar to Denmark, the reform would mean an increase in labour market flexibility. Less straightforward is the case of countries with very 'flexible' regimes, where the parallel moves toward less flexibility and more generosity work in opposite directions.

The estimated changes in the short-run elasticity resulting from the policy shift described above are reported in Figure 3, in descending order, together with the 95% confidence intervals. The computation was based on expression (4) and the estimated coefficients in Table A.9, column 7, 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 3: Impact of a flexicurity reform on the 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. The highest increases of above 0.2 are obtained for Great Britain, Germany, Australia and Japan. Positive and signficant effects are obtained also for Italy, Norway, New Zealand, Sweden, the Netherlands, Finland, Austria and Switzerland. At the other end, the estimated effects are small or not statistically significant for Portugal, Canada, Ireland, France, Belgium, the US and Greece.

The group of countries with the highest expected effects includes countries that have considerably less generous systems than Denmark. 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 a statistically signifiant positive impact. Under a flexicurity reform, these two countries would need to reduce the flexibility of their labour markets, which would counteract other elements of the reform package.

These results indirectly lends 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 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 hightened the emphasis on *flexicurity* reforms. This paper examines the responsiveness of employment to output fluctuations in a panel of 20 OECD countries, paying particular attention to the role of the key pillars of the flexicurity as well as their interactive effects in driving this outcome.

Interesting results emerge. We find that, for an average country, 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.

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Appendix: Figures and Tables

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

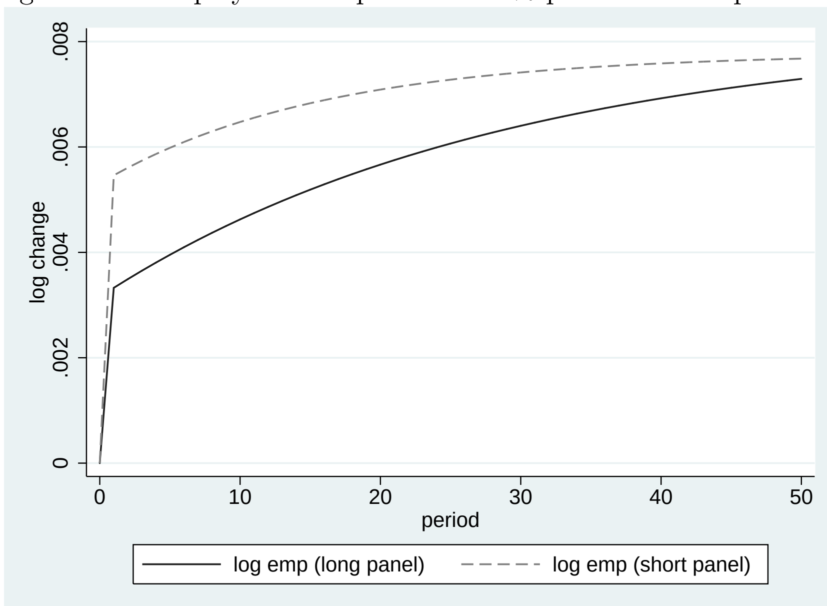

Figure A.2: Benefit replacement rate by country

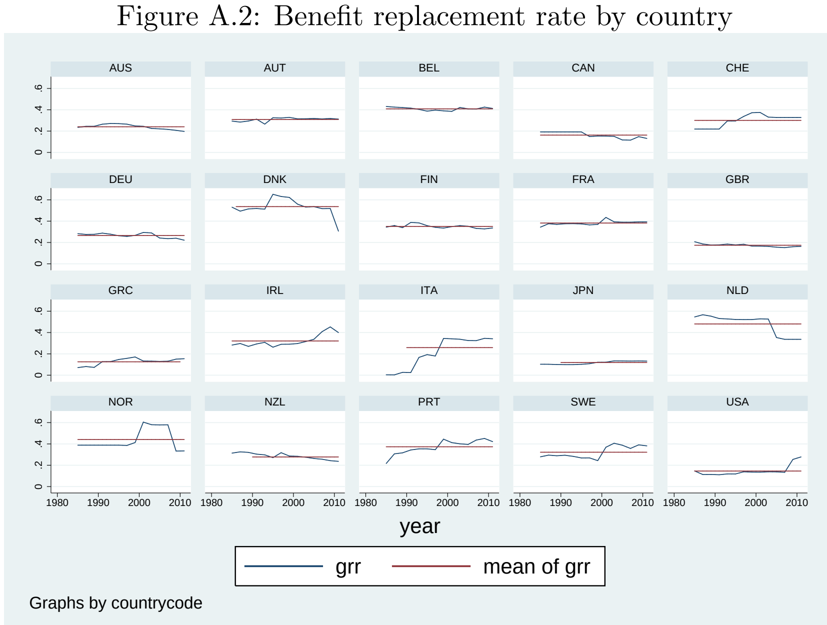

Figure A.3: Labour market flexibility by country

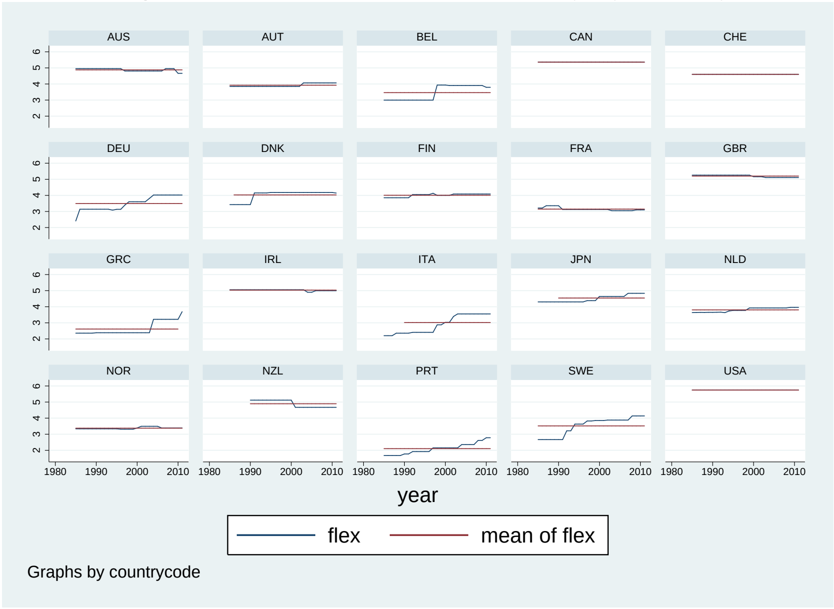

Figure A.4: Training expenditures per unemployed by country

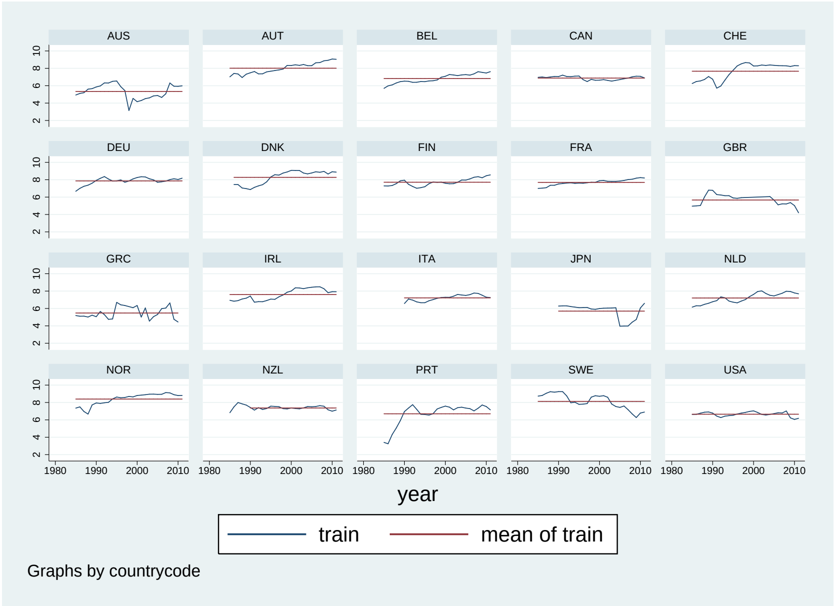

Figure A.5: Elasticity and the policy variables

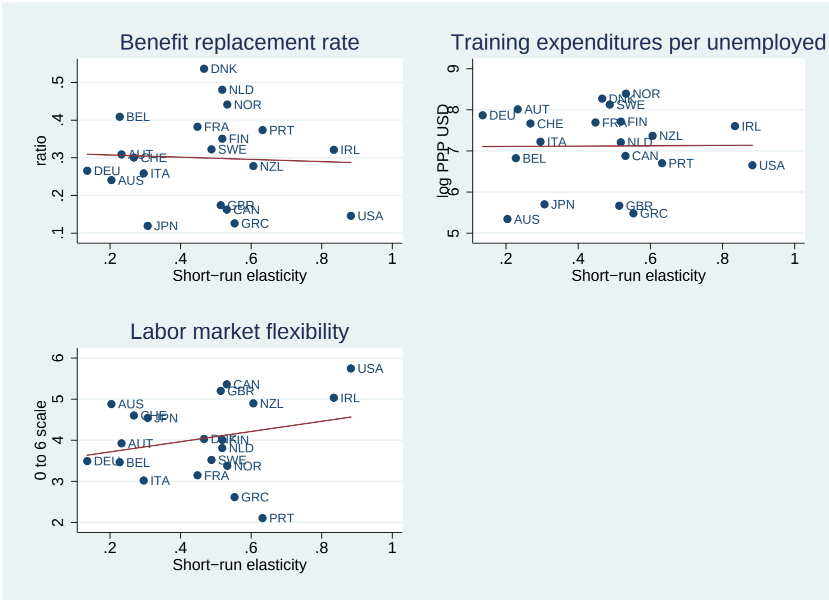

Table A.1: Variable description

Variable	Description	Data Source
	Core variables	
employment	Number of persons engaged (millions)	Penn World Table 9.0[a]
output	Real GDP at constant national prices (mil. 2011US$)	Penn World Table 9.0
	Policy variables	
grr	Gross unemployment benefit replacement rate[b]	OECD
train	Training expenditures per unemployed person (PPP US$, log)	OECD
flex	Labour market flexibility indicator (0-6 scale)	OECD, EPL indicators
	Control variables	
openness	Trade openness (export+import)/GDP in real 2010US$	World Bank, WDI
uniondensity	Trade union density (share)	OECD
govsize	Government consumption as share of GDP	OECD, National Accounts

Notes: [a] The Penn World Table 9.0 is described in Feenstra, Inklaar and Timmer (2015). [b] The grr series is available for odd years only and subject to a methodological change in 2005. It was calibrated for the average production worker until 2005 and for the average worker afterwards. To produce a full time series, we fill the even years with linear interpolation. To account for the break in 2005, we link the annual changes in the new grr series for years after 2005 to the level of the old one.

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	-1.444	0.453	0.238	0.169	0.137^b	-2.438	0.479	0.259	0.187	0.152
AUT	-0.182	1.070	0.564	0.394	0.311	-1.125	1.023	0.542	0.379	0.298
BEL	0.499	1.201	0.618	0.424	0.328	-2.220	0.983	0.525	0.369	0.292
CAN	-1.639	1.041	0.544	0.380	0.299	-2.206	0.979	0.520	0.367	0.291
CHE	-0.959	0.576	0.307	0.219	0.177	-3.978^a	0.427	0.238	0.174	0.144^b
DEU	-0.516	0.443	0.245	0.183	0.155	-1.748	1.095	0.585	0.412	0.326
DNK	-2.804	0.494	0.274	0.204	0.171	-1.953	0.848	0.463	0.332	0.267
FIN	-1.175	0.431	0.224	0.158	0.127^b	-0.949	0.801	0.435	0.313	0.252
FRA	-1.621	0.511	0.269	0.191	0.153	-2.444	1.098	0.582	0.408	0.322
GBR	-0.748	1.038	0.547	0.386	0.307	-2.006	0.217	0.121^b	0.092^b	0.078^b
GRC	-0.277	0.379	0.207	0.153	0.128^b	-1.415	0.822	0.441	0.312	0.250
IRL	-1.418	0.984	0.504	0.345	0.267	-0.944	0.584	0.301	0.208	0.163
ITA	-2.052	0.302	0.163	0.118^b	0.098^b	-0.684	1.212	0.642	0.448	0.350
JPN	-0.789	1.185	0.617	0.427	0.333	-2.783	1.231	0.647	0.449	0.350
NLD	-0.863	0.785	0.408	0.284	0.224	-1.154	0.719	0.383	0.272	0.217
NOR	-1.729	0.273	0.145^b	0.106^b	0.088^b	-0.014	1.114	0.589	0.412	0.324
NZL	-1.224	0.593	0.306	0.211	0.165	-2.471	0.367	0.201	0.146	0.121^b
PRT	1.556	0.793	0.425	0.303	0.243	-0.206	1.067	0.568	0.401	0.317
SWE	-1.572	0.631	0.331	0.233	0.186	-3.149	0.472	0.259	0.188	0.154
USA	-0.152	1.116	0.586	0.411	0.325	-1.278	0.617	0.339	0.248	0.205

Notes: ADF: Augmented Dickey-Fuller test. The test regression is specified to include a constant and a time trend and no lagged dependent variable. H0: series has a unit root (against being trend stationary). ADF critical values: 1% -4.141, 5% -3.496, 10% -3.178. 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. ^{a} ADF rejects unit root against trend stationarity at 5% critical value. ^{b} KPSS cannot reject stationarity at 5% critical value.

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	-4.102	0.150^d	0.099	0.086	0.084	-5.654	0.125	0.107	0.103	0.102
AUT	-3.185	0.097	0.067	0.058	0.054	-5.170	0.085	0.080	0.082	0.077
BEL	-4.151	0.257^d	0.182^d	0.157^d	0.144	-5.189	0.138	0.132	0.124	0.118
CAN	-4.325	0.100	0.072	0.066	0.069	-4.722	0.111	0.092	0.089	0.092
CHE	-3.963	0.163^d	0.103	0.087	0.081	-4.826	0.207^d	0.147^d	0.133	0.126
DEU	-4.043	0.113	0.076	0.069	0.070	-5.310	0.044	0.039	0.047	0.057
DNK	-4.593	0.087	0.063	0.061	0.063	-5.366	0.096	0.088	0.090	0.090
FIN	-3.286	0.138	0.084	0.069	0.064	-4.547	0.071	0.052	0.050	0.051
FRA	-4.160	0.103	0.069	0.062	0.060	-3.305	0.188^d	0.146^d	0.135	0.128
GBR	-3.628	0.072	0.047	0.042	0.045	-5.520	0.058	0.047	0.049	0.054
GRC	-3.591	0.396^d	0.249^d	0.197^d	0.174^d	-4.053	0.209^d	0.156^d	0.130	0.113
IRL	-3.423	0.286^d	0.176^d	0.136	0.117	-3.409	0.252^d	0.155^d	0.125	0.110
ITA	-3.894	0.213^d	0.139	0.117	0.104	-4.539	0.045	0.047	0.054	0.055
JPN	-3.406	0.096	0.065	0.056	0.052	-3.347	0.174^d	0.141	0.126	0.113
NLD	-3.307	0.273^d	0.168^d	0.141	0.135	-3.620	0.179^d	0.126	0.111	0.102
NOR	-3.493	0.089	0.055	0.048	0.048	-4.213	0.067	0.051	0.046	0.048
NZL	-4.473	0.266^d	0.187^d	0.155^d	0.138	-5.851	0.098	0.082	0.079	0.076
PRT	-4.115	0.395^d	0.273^d	0.234^d	0.206^d	-4.173	0.068	0.055	0.055	0.054
SWE	-3.968	0.147^d	0.097	0.088	0.089	-5.346	0.156^d	0.125	0.126	0.124
USA	-4.521	0.075	0.056	0.055	0.061	-5.030	0.072	0.058	0.058	0.063

Notes: ADF: Augmented Dickey-Fuller test. The test regression is specified to include a constant and no lagged dependent variable. H0: series has a unit root. ADF critical values:1% -3.576 5% -2.928 10% -2.599. 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. ^{c} ADF cannot reject the unit root at 5% critical value. ^{d} KPSS rejects stationarity at 5% critical value.

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*	-1.103	0.135	-1.208	0.114
Fisher-ADF				
Ho: All panels contain unit roots; Ha: At least one panel is stationary				
Inverse chi-squared(40) P	33.967	0.738	39.740	0.482
Inverse normal Z	1.582	0.943	1.006	0.843
Inverse logit t(104) L*	1.710	0.955	1.341	0.909
Modified inv. chi-squared Pm	-0.675	0.750	-0.029	0.512
Hadri				
Ho: All panels are stationary; Ha: Some panels contain unit roots				
z	66.639	0.000	70.487	0.000

Notes: 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*	-10.516	0.000	-12.125	0.000
Fisher-ADF				
Ho: All panels contain unit roots; Ha: At least one panel is stationary				
Inverse chi-squared(40) P	163.381	0.000	175.296	0.000
Inverse normal Z	-9.115	0.000	-9.563	0.000
Inverse logit t(104) L*	-9.990	0.000	-10.701	0.000
Modified inv. chi-squared Pm	13.794	0.000	15.127	0.000
Hadri				
Ho: All panels are stationary; Ha: Some panels contain unit roots				
z	13.754	0.000	14.606	0.000

Notes: 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

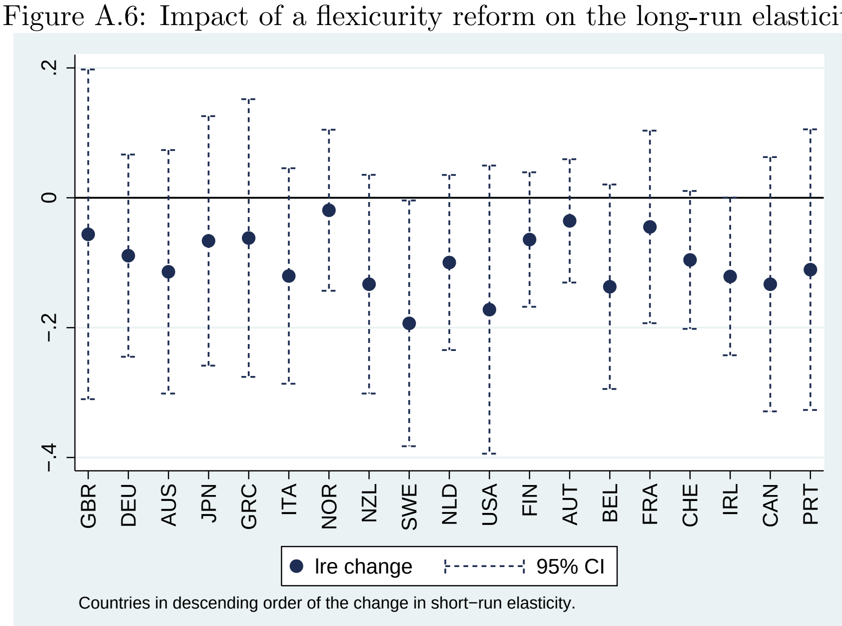

Table A.6: Johansen rank tests for cointegration

Country code (ISO)	Max rank	Trace statistics	Max-eigenvalue statistics	Number of CI[1]	Number of CI[2]
AUS	0	24.802	17.990	0	1
	1	6.812	6.812		
AUT	0	44.055	37.558	1	1
	1	6.498	6.498		
BEL	0	53.627	43.557	1	1
	1	10.070	10.070		
CAN	0	30.974	22.662	1	1
	1	8.312	8.312		
CHE	0	37.276	26.355	1	1
	1	10.921	10.921		
DEU	0	30.942	20.660	1	1
	1	10.282	10.282		
DNK	0	37.912	28.236	1	1
	1	9.676	9.676		
FIN	0	33.554	25.769	1	1
	1	7.786	7.786		
FRA	0	42.375	33.833	1	1
	1	8.542	8.542		
GBR	0	23.832	15.354	0	1
	1	8.478	8.478		
GRC	0	26.543	19.123	1	1
	1	7.419	7.419		
IRL	0	15.346	9.065	0	0
	1	6.281	6.281		
ITA	0	44.392	32.466	1	1
	1	11.926	11.926		
JPN	0	31.158	25.620	1	1
	1	5.539	5.539		
NLD	0	36.216	29.966	1	1
	1	6.250	6.250		
NOR	0	35.986	25.112	1	1
	1	10.874	10.874		
NZL	0	19.631	15.798	0	1
	1	3.833	3.833		
PRT	0	27.173	16.274	1	1
	1	10.899	10.899		
SWE	0	24.650	19.768	0	1
	1	4.882	4.882		
USA	0	13.897	12.392	0	1
	1	1.505	1.505		
5% critical value	0	25.32	18.96		
	1	12.25	12.52		

Notes: Results from Johansen's cointegrating rank test, 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: Cross-country correlation of policies

	(grr,train)	(grr,flex)	(flex,train)
Full period			
1985-2011	0.641***	-0.343	-0.094
Sub-periods			
1985-1994	0.354	-0.187	-0.021
1995-2004	0.638***	-0.326	-0.122
2005-2011	0.772***	-0.532**	-0.325

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

Table A.8: Estimates with policies (first-difference specification)

Depvar: Δe	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Δy	0.523***	0.532***	0.527***	0.530***	0.483***	0.566***	0.517***
	(0.046)	(0.048)	(0.047)	(0.045)	(0.047)	(0.040)	(0.040)
Δy x $\tilde{\text{grr}}_{-1}$	0.502*			0.869**	0.760**	0.695**	0.690**
	(0.283)			(0.340)	(0.345)	(0.295)	(0.312)
Δy x $\tilde{\text{train}}_{-1}$		0.022		-0.036	0.011	-0.003	0.031
		(0.028)		(0.034)	(0.034)	(0.032)	(0.030)
Δy x $\tilde{\text{flex}}_{-1}$			0.068**	0.080**	0.055	0.093***	0.067*
			(0.033)	(0.033)	(0.040)	(0.029)	(0.035)
Δy x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$					0.270		0.245
					(0.201)		(0.186)
Δy x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.956***		-0.913***
					(0.369)		(0.330)
Δy x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					0.090**		0.062**
					(0.036)		(0.031)
Δy x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.026		-0.229
					(0.292)		(0.261)
$\tilde{\text{grr}}_{-1}$	-0.027**			-0.037**	-0.038**	-0.026**	-0.031**
	(0.014)			(0.015)	(0.015)	(0.013)	(0.013)
$\tilde{\text{train}}_{-1}$		0.001		0.002	0.000	0.001	-0.001
		(0.001)		(0.001)	(0.001)	(0.001)	(0.001)
$\tilde{\text{flex}}_{-1}$			-0.001	-0.001	0.001	-0.003	-0.001
			(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
$\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$					-0.012		-0.014*
					(0.009)		(0.008)
$\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.012		-0.014
					(0.014)		(0.013)
$\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.004**		-0.004***
					(0.002)		(0.001)
$\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.027**		-0.025**
					(0.012)		(0.012)
Control variables						✓	✓
Country dummies	✓	✓	✓	✓	✓	✓	✓
Year dummies	✓	✓	✓	✓	✓	✓	✓
Observations	516	516	516	516	516	516	516
R-squared	0.652	0.648	0.652	0.664	0.686	0.714	0.737

Notes: Estimation of (3) with OLS. 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. Controls variables include the annual changes in trade openness, trade union density and government size. *** p<0.01, ** p<0.05, * p<0.1.

Table A.9: Estimates with policies (ecm specification)

Depvar: Δe	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Δy	0.510***	0.531***	0.523***	0.520***	0.451***	0.569***	0.495***
	(0.042)	(0.047)	(0.045)	(0.042)	(0.046)	(0.038)	(0.042)
Δy x $\tilde{\text{grr}}_{-1}$	0.480*			0.852***	0.696**	0.751***	0.649**
	(0.270)			(0.322)	(0.318)	(0.273)	(0.292)
Δy x $\tilde{\text{train}}_{-1}$		0.017		-0.032	0.012	0.000	0.037
		(0.027)		(0.032)	(0.032)	(0.030)	(0.030)
Δy x $\tilde{\text{flex}}_{-1}$			0.076**	0.081**	0.083*	0.089***	0.091**
			(0.034)	(0.033)	(0.043)	(0.029)	(0.038)
Δy x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$					0.299		0.301*
					(0.194)		(0.180)
Δy x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.783**		-0.779**
					(0.376)		(0.338)
Δy x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					0.053		0.036
					(0.043)		(0.038)
Δy x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.244		-0.364
					(0.311)		(0.277)
e_{-1}	-0.070***	-0.071***	-0.075***	-0.084***	-0.100***	-0.074***	-0.078***
	(0.013)	(0.014)	(0.014)	(0.015)	(0.016)	(0.014)	(0.015)
e_{-1} x $\tilde{\text{grr}}_{-1}$	-0.020			-0.017	0.064	-0.042	0.042
	(0.030)			(0.032)	(0.058)	(0.031)	(0.054)
e_{-1} x $\tilde{\text{train}}_{-1}$		0.000		0.001	0.002	0.005	0.001
		(0.003)		(0.003)	(0.006)	(0.003)	(0.005)
e_{-1} x $\tilde{\text{flex}}_{-1}$			0.003	0.004	-0.011	0.009*	-0.005
			(0.005)	(0.005)	(0.008)	(0.005)	(0.007)
e_{-1} x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$					0.033		0.013
					(0.048)		(0.046)
e_{-1} x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.030		-0.016
					(0.062)		(0.059)
e_{-1} x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					0.009*		0.008*
					(0.005)		(0.004)
e_{-1} x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					0.136***		0.098**
					(0.049)		(0.046)

Table continues on the next page.

Depvar: Δe	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Continued from the previous page.							
y_{-1}	0.055***	0.052***	0.052***	0.060***	0.074***	0.046***	0.054***
	(0.012)	(0.014)	(0.013)	(0.012)	(0.014)	(0.011)	(0.012)
y_{-1} x $\tilde{\text{grr}}_{-1}$	0.042			0.038	-0.028	0.061**	-0.008
	(0.028)			(0.030)	(0.050)	(0.028)	(0.046)
y_{-1} x $\tilde{\text{train}}_{-1}$		0.000		-0.001	-0.005	-0.005	-0.004
		(0.002)		(0.003)	(0.005)	(0.003)	(0.005)
y_{-1} x $\tilde{\text{flex}}_{-1}$			0.000	-0.001	0.010	-0.003	0.006
			(0.004)	(0.004)	(0.007)	(0.004)	(0.006)
y_{-1} x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$					-0.067		-0.041
					(0.049)		(0.047)
y_{-1} x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					0.007		0.002
					(0.057)		(0.054)
y_{-1} x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.006		-0.004
					(0.004)		(0.004)
y_{-1} x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.110***		-0.071*
					(0.042)		(0.039)
$\tilde{\text{grr}}_{-1}$	-0.552*			-0.514	0.193	-0.767**	-0.022
	(0.316)			(0.333)	(0.550)	(0.315)	(0.506)
$\tilde{\text{train}}_{-1}$		-0.001		0.019	0.065	0.053	0.051
		(0.027)		(0.034)	(0.061)	(0.033)	(0.055)
$\tilde{\text{flex}}_{-1}$			-0.015	-0.002	-0.118	0.021	-0.077
			(0.047)	(0.049)	(0.073)	(0.045)	(0.071)
$\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$					0.803		0.500
					(0.551)		(0.527)
$\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					-0.046		-0.005
					(0.631)		(0.595)
$\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					0.057		0.040
					(0.049)		(0.047)
$\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$					1.175**		0.742*
					(0.461)		(0.431)
Control variables						✓	✓
Country dummies	✓	✓	✓	✓	✓	✓	✓
Year dummies	✓	✓	✓	✓	✓	✓	✓
Observations	516	516	516	516	516	516	516
R-squared	0.680	0.670	0.676	0.697	0.732	0.740	0.767

Notes: OLS estimation of (3) extended with interactions of e_{-1} and y_{-1} with the policy variable(s). 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. Controls variables include the annual changes in trade openness, trade union density and government size. *** p<0.01, ** p<0.05, * p<0.1.

Table A.10: Effects on the long-run elasticity of σ-increases in the policy variables

Policy variable	without interactions	with interactions average	flexible	generous
grr	0.055***	0.036	0.047	0.009
	(0.019)	(0.029)	(0.032)	(0.032)
train	-0.028	-0.049*	0.021	-0.078*
	(0.020)	(0.028)	(0.067)	(0.041)
flex	0.032	0.032	0.061	0.031
	(0.025)	(0.026)	(0.045)	(0.039)

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 (6), those with interactions on column (7) of Table A.9. Average, flexible and generous regimes differ in the values the non-changing policy variables are assumed to take. Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A.11: Policy distance from Denmark in 2010

Country code (ISO)	Policy variable grr	train	flex
AUS	-0.210	-2.988	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	-0.891	-0.161
DNK	0.000	0.000	0.000
FIN	-0.081	-0.461	-0.097
FRA	-0.019	-0.674	-1.083
GBR	-0.251	-3.892	0.926
GRC	-0.260	-4.470	-0.962
IRL	0.014	-0.989	0.817
ITA	-0.069	-1.623	-0.626
JPN	-0.279	-2.861	0.655
NLD	-0.076	-1.129	-0.218
NOR	-0.078	-0.120	-0.793
NZL	-0.172	-2.010	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.