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

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

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**Abstract**

Labour market reforms in the direction of 'flexicurity' have been widely endorsed as a means to increasing 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 panel of OECD countries, which also incorporates the period of the Great Recession, and distinguish between passive and active labour market policy types. We find that the effects of any single policy change are shaped by the broader existing policy-mix within which it takes place.

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

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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 these inter-country variations 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. 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. 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.[2] 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 *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

[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.*, 2020), 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.*, 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 *et al.*, 2014).

been forcefully demonstrated theoretically – e.g. Davoine and Keuschnigg (2015), Dabusinskas *et al.* (2016) and Molana *et al.* (2020) – 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.

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. The long-run elasticity, however, has 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 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*, *flexible* and *generous*. 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.* (2018, 2020), 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 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., 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) 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 cycles, 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 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) 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

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 (e.g., Hendry, 1995). An advantage of this 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.[3] 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 *et al.* (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 ($y$) 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

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

$t$ 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 bracket) and a random error term ($\epsilon_{i,t}$) capturing contemporaneous employment shocks. The term in the bracket is the long-run equilibrium relationship between employment and output, where $\beta$ is assumed to be negative and $-\beta$ measures the long-run employment to output elasticity. The long-run relationship also allows for a deterministic linear time trend, which may represent long-run technological or demographic changes that are common across countries. The short-run employment-output elasticity is captured by $\gamma$.

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 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 $\alpha$.

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.

We reformulate equation (1) into the following estimating 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 $\alpha$ remains the speed of adjustment parameter, $\gamma$ is the short-run elasticity, while the long-run elasticity is obtained as $-\frac{\beta^*}{\alpha}$. Estimating this equation instead of (1) allows for more flexibility in the treatment of the common time trend, as it assumes year-specific fixed effects ($\delta_t$) instead of a linear time trend. Moreover, the equation accounts for country-specific constants ($\delta_i$). As a robustness check, we will also experiment with adding country-specific linear time trends, which can account for the possibility that secular trends, e.g. in terms of demography or technology, differ from country to country.

In the above framework, it is possible to assess the existence of a cointegrating relationship by testing whether $\alpha$ is significantly smaller than zero (Ericsson and MacKinnon, 2002). Should the estimated value of $\alpha$ be non-negative, then the cointegrating relationship is considered as nonexistent, 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 $\gamma$. 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 *et al.*, 2012; Bassanini, 2012) and include interactions of the right-hand side variables with the policy variables in the estimating 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 the ease of interpretation, the policy variables are 'centered' within each year, i.e. expressed as deviations from their country means as $\tilde{P}_{i,t} = P_{i,t} - \frac{1}{N}\sum_{j=1}^{N} P_{j,t}$. 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,

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

where $k$ is the index for the different policy measures and – possibly – their interactions. The estimate for $\gamma_1$ will always give the short-run elasticity for the average country.[4] The estimated $\gamma_2$ 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,

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

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

Policy variables can, of course, relate endogeneously 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. 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., union density, openness to international trade or the size of the government, which 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 the employment responsiveness (e.g., Duval *et al.*, 2020; Ciminelli *et al.*, 2018). 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* instead 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 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 employ-

[5]The Penn World Table we use is described in Feenstra *et al.* (2015). The data is available for download at `www.ggdc.net/pwt`.

ment and output. 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 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 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 differences.

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 exists, which we test in two ways. 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 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 examinig the statistical significance of the speed of adjustment parameter when estimating equation (2) 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.[6]

# 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 (2) and (3) 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 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 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 estimate of the long-run employment-output elasticity, obtained using error correction specification, is around 0.8 for the average country on both samples. This means that if output permanently increases by 1%, employment will also rise by 0.8% in the long run, ceteris paribus. 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 instantaneous 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.[7]

[6] 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.

[7] The response function on Figure A.1 is produced by making dynamic forecasts of our estimated model with and

Table 1: Elasticity estimates

| Depvar: $\Delta e$ | 1960-2014 | | 1986-2012 | |
|---|---|---|---|---|
| $\Delta y$ (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 models (2) and (3) are based on OLS. 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.

Our elasticity estimates are roughly consistent with those found for developed countries in previous literature (e.g. Kapsos, 2006; Crivelli *et al.*, 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.

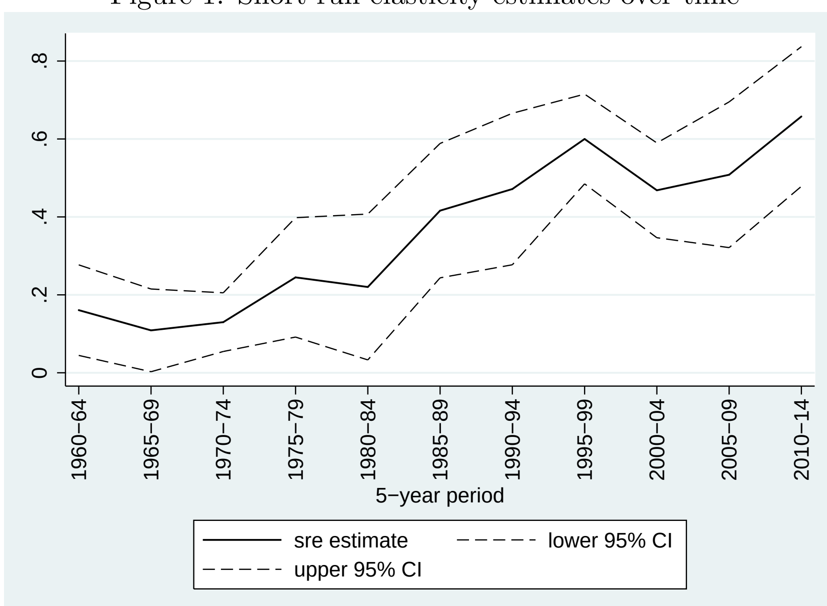
A robustness exercise is reported in Table A.7 which shows that the short-run elasticity estimates are remarkably robust to the inclusion of country-specific linear time trends. We also find evidence that the cointegrating relationship exists under the assumption of such trends and that - at least on the long panel - the speed of adjustment parameter and the long-run elasticity do not change significantly compared to the baseline estimates. Nevertheless, on the short panel, the latter two estimates 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 specific trends in the error correction (EC) regression on the short panel carries the risk of model overfitting. For this reason, in what follows we only

without a 1% shock to output and taking the difference of the two forecast paths.

include country specific trends in the more parsimonious first differenced (FD) specification.

Next, we investigate more closely how the short-run elasticity changes over time. To this end, we interact the output change variable with (five-year) period dummies when estimating equation (2) on the long panel. The resulting period-specific elasticity estimates are plotted in Figure 1.[8] 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 – for instance, as suggested by Blanchard *et al.* (2006) it 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.

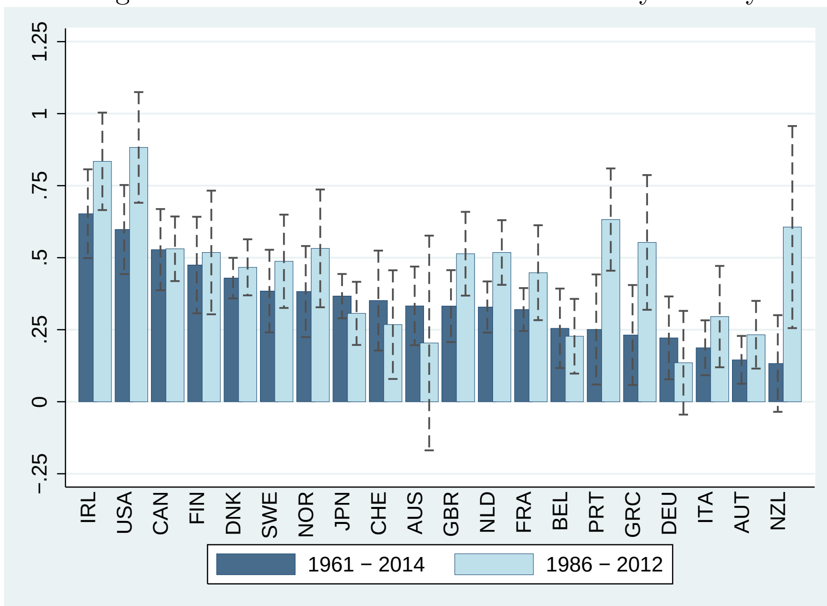
Figure 1: Short-run elasticity estimates over time




In order to establish the extent of inter-country heterogeneity in employment responsiveness, we estimate model (2) by interacting the output change variable with country dummies to obtain country-specific elasticities. Figure 2 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. It is apparent that considerable inter-country variations exist in the responsiveness of employment to output. Countries with the largest elastic-

[8]Estimating the FD specification (3) or including country-specific trends in the regression result in very similar graphs.

Figure 2: Estimated short-run elasticities by country




ities – such as the US and Ireland on the left-hand side of the graph – have at least three times higher employment responsiveness than countries on the right-hand side (e.g. Austria, Germany). The estimates for some countries also change considerably when the sample is shortened to the more recent decades. In particular, the elasticities for Greece, New Zealand and Portugal, which all have below-average elasticities in the long panel, increase to above average.[9]

In the remaining part of this paper, we focus on the role of labour market institutions and policies in driving inter-country heterogenous responses in employment-output elasticities.

## 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 considered as 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 needs for the flexibility required to adjust to negative shocks with that for adequate

[9]These findings are largely unchanged (though with some changes in the country ranking) when estimating the first-difference specification or including country-specific trends in the regression.

social safety nets (see, e.g., Blanchard *et al.*, 2014). The descriptive statistics of the three policy variables are reported in Table 2 below.

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.[10] *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.[11]

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.2). 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 in both hiring and firing is empirically challenging. However, arguably – as is well demonstrated in the theoretical literature (see, for

[10] 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.

[11] 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.

instance, Zanetti, 2011; Cacciatore and Fiori, 2016) – the 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.[12] Our *flex* index measures, on a 0 to 6 scale, how easy it is for employers to lay off workers (firing) or to employ them on temporary work contracts (hiring), as laid down by a country's regulatory system.[13] Hence, a higher value of *flex* means a more flexible labour market, which entail both lower costs of hiring and firing. 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 *targeted* 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 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

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

[13] Following the OECD practice (OECD, 2013), we take the weighted average of the employment protection subindices 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* variable.

suggests the relatively 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.8 reveal that countries with generous unemployment benefits tend to 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 (Blanchard *et al.*, 2014; 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).

Thus, we estimate the EC equation while introducing the policy variables and their interactions with the right-hand-side variables and report the estimation results in Table A.10. The coefficient estimates in the different columns of the table are reasonably stable, regardless whether policies are included individually or jointly and whether policy interactions are also accounted for. This stability also holds for the inclusion of country-time-specific control variables, which is shown in the last two columns. The set of control variables includes the size of the working-age population, openness to international trade, trade union density, and the size of the government (see a more detailed description in Table A.1), that can potentially affect employment elasticities. Panel unit root tests indicate that all of these variables are first-order integrated except for the working-age population, which is second-order integrated (Table A.9). This motivates us to include in the regression the annual changes of trade openness, trade union density and government size and the change in the annual change (second-order difference) of the working-age population. All but trade openness turn out to be significant predictors of employment growth. At the same time, their inclusion leaves the rest of the coefficient estimates roughly unchanged.

The estimates in Table A.10 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 figure of 0.701 in the second row of column 6, 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 0.07 higher-than-average short-run elasticity. Similarly, a one-point higher flexibility index corresponds to a 0.08 higher elasticity. 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.

Columns 5 and 7 consider all possible interactions between the policy variables. 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 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.

The above results are largely unchanged when we estimate the FD specification (Table A.11) or the FD specification with country-specific linear time trends (Table A.12). A minor difference is that in the FD regression the interaction of *train* and *flex* is significantly positively associated with the short-run employment to output elasticity. This finding, however, disappears when country-specific time trends are also included.

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 *et al.*, 2018; Shimer, 2005; Zanetti, 2011) as our estimates reveal. 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 job 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 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 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 greater 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). The offsetting effects of unemployment generosity and flexibility produces 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 (e.g. Onwordi, 2016; Pissarides, 2009; Stähler and Thomas, 2012), a positive shock to productivity is likely to result in greater elasticity in the short-run since it reduces the profitability of job matches. However, as argued by Molana *et al.* (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 supports 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*, *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 (5) to calculate the effects on the short-run elasticity, using the estimated coefficients in Table A.10, reported either in column 6 (without policy interactions) or in column 7 (with policy interactions). 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 $\hat{\gamma}_2^{\tilde{\text{grr}}} \times \hat{\sigma}^{\text{grr}} = 0.701 \times 0.125 = 0.088$, where $\hat{\gamma}_2^{\tilde{\text{grr}}}$ is the estimated coefficient for the interaction of $\Delta y$ with the gross replacement rate in column 6, 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 $\Delta y$ in column 7 of Table A.10 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.

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 0.088 and 0.078, respectively). 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 *flexible* regimes, making the unemployment benefit more

Table 3: Effects on the short-run elasticity of $\sigma$-increases in the policy variables

| Policy variable | without interactions | with interactions average | flexible | generous |
|---|---|---|---|---|
| grr | 0.088*** | 0.077** | 0.004 | 0.197*** |
| | (0.034) | (0.036) | (0.055) | (0.058) |
| train | 0.003 | 0.039 | 0.102 | 0.066 |
| | (0.033) | (0.033) | (0.137) | (0.043) |
| flex | 0.078*** | 0.083** | 0.122*** | 0.017 |
| | (0.029) | (0.037) | (0.038) | (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. Estimates without interactions are based on column (6), those with interactions on column (7) of Table A.10. *Average*, *flexible* and *generous* 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.

generous does not result in 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 a *generous* 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 *et al.* (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. As evident in column 2 (where complementarities are not taken into account), however, implementing each of these policies in an average regime, will have significant effects. 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 robust to changing the model specification into FD with or without country-specific trends (Table A.14).

Finally, we calculate the corresponding changes in the long-run elasticity, which we report in Table A.13.[14] The numbers suggest that the effects of these policies found on the short-run elasticity

[14]This we do by evaluating expression (6) before and after the one-stdev 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.

mostly fade out in the long run. Training is the only policy with a (weakly) significant and negative long-run impact on the employment elasticities 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.*, 2018).[15]

## 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-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.15 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 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, 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 3, in descending order, together with the 95% confidence intervals. The computation was based on expression (5) and the estimated coefficients in Table A.10, column 7,

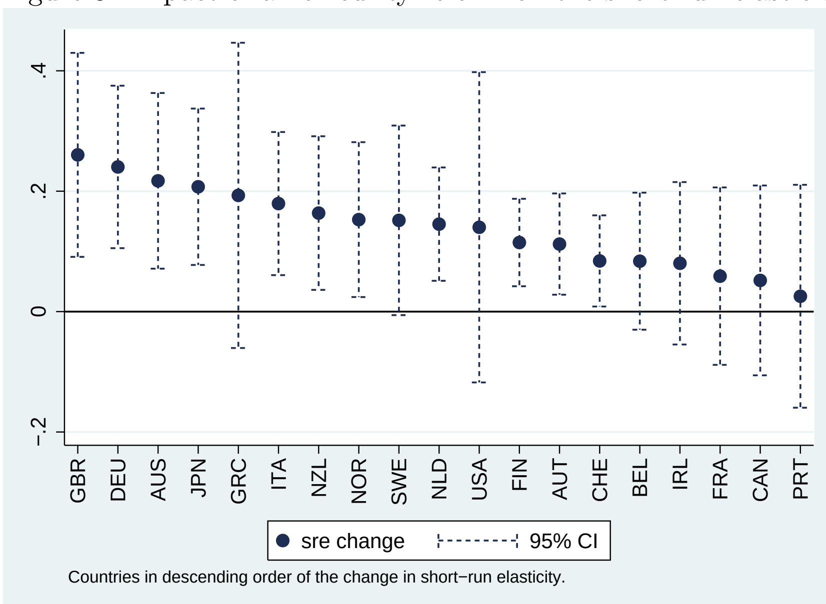
[15] In a search and matching open economy model, Molana *et al.* (2018) show that increases in training expenditure can reduce employment volatility in response to exogenous shocks even in the presence of more generous unemployment insurance.

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 significant effects are obtained also for Italy, New Zealand, Norway, the Netherlands, Finland, Austria and Switzerland. At the other end, the estimated effects are small or not statistically significant for Portugal, Canada, France, Ireland, Belgium, the US and Greece.

The group of countries with the highest expected effects includes those with 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 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, either with or without country trends. Table A.16 compares the baseline estimates with the results of the two robustness checks. The groups of countries for which statistically significant increases in the short-run elasticity are predicted are almost identical under the three specifications, and the relative sizes of the country estimates also remain largely unchanged.

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 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

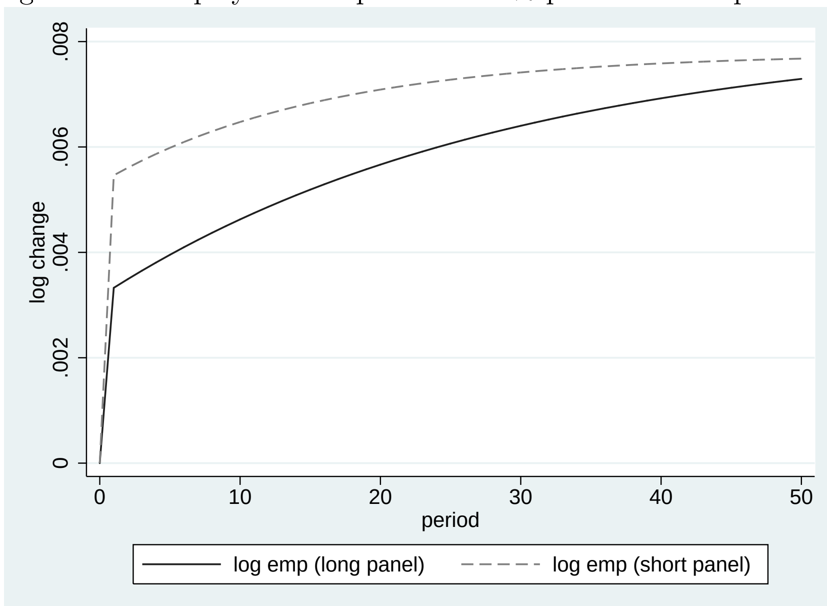


Figure A.2: Benefit replacement rate by country

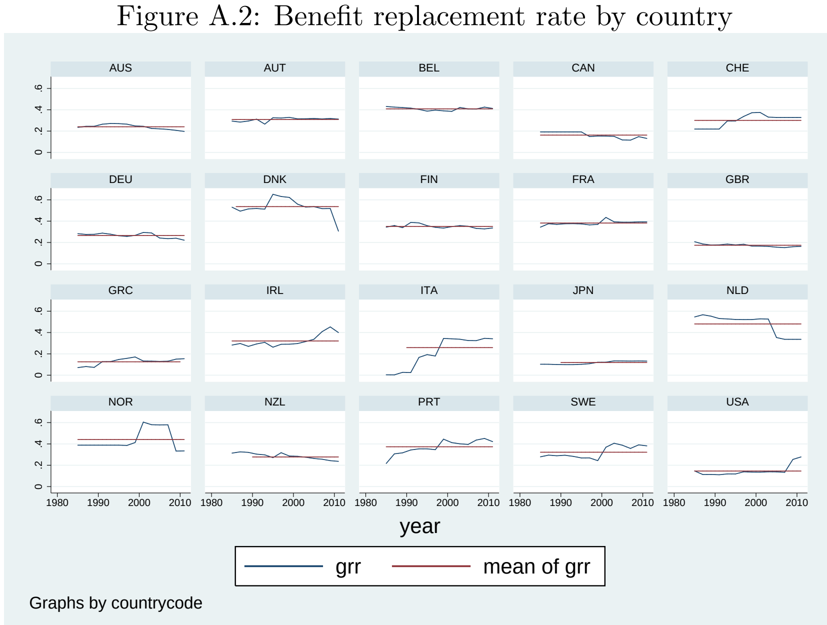

Table A.1: Variable description

| Variable | Description | Data Source |
|---|---|---|
| | *Core variables* | |
| e (employment) | Number of persons engaged (millions, in log) | Penn World Table 9.0 |
| y (output) | Real GDP at constant national prices (mil. 2011US$, in log) | Penn World Table 9.0 |
| | *Policy variables* | |
| grr | Gross unemployment benefit replacement rate | OECD |
| train | Training expenditures per unemployed person (PPP US$, log) | OECD |
| flex | Labour market flexibility indicator (0-6 scale) | OECD, EPL indicators |
| | *Control variables* | |
| wapop | Working-age population (ages 20-64, mil. persons, in log) | OECD |
| openness | Trade openness (export+import)/GDP in real 2010US$ | World Bank, WDI |
| union | Trade union density (share) | OECD |
| govsize | Government consumption as share of GDP | OECD, National Accounts |

Notes: The Penn World Table 9.0 is described in Feenstra *et al.* (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 | -1.444 | 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 code (ISO) | dlog employment ($\Delta e$) ADF | KPSS0 | KPSS1 | KPSS2 | KPSS3 | dlog output ($\Delta y$) 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 ($e$) Statistic | p-value | log output ($y$) 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.

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

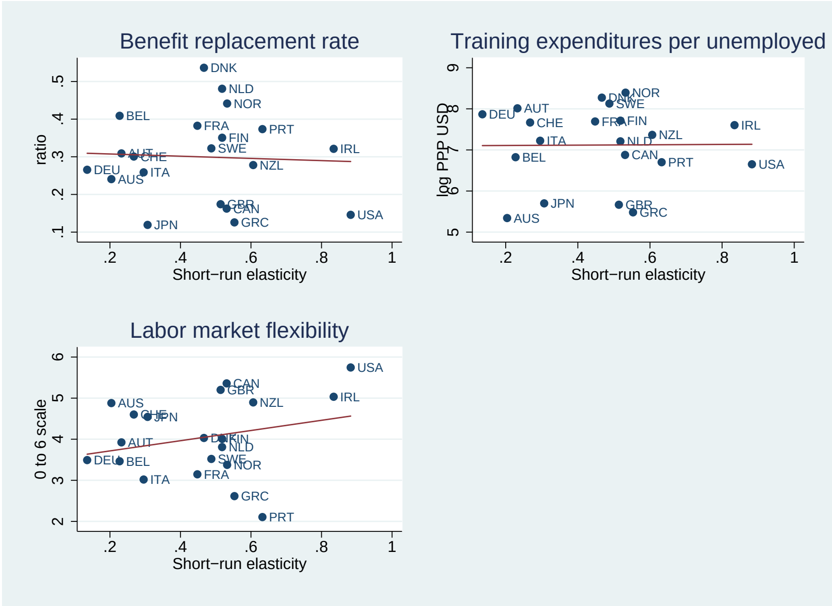


Table A.5: Panel unit root tests for first differences

| | dlog employment ($\Delta e$) | | dlog output ($\Delta y$) | |
|---|---|---|---|---|
| | 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.

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: Robustness: Elasticity estimates with country trends

| Depvar: $\Delta e$ | 1960-2014 | | 1986-2012 | |
|---|---|---|---|---|
| $\Delta y$ (Short-run elasticity) | 0.325*** | 0.345*** | 0.509*** | 0.535*** |
| | (0.034) | (0.032) | (0.040) | (0.035) |
| $e_{-1}$ (Speed of adjustment) | | -0.047*** | | -0.151*** |
| | | (0.011) | | (0.022) |
| | | [-4.24] | | [-6.76] |
| $y_{-1}$ | | 0.058*** | | 0.222*** |
| | | (0.008) | | (0.022) |
| Country dummies | ✓ | ✓ | ✓ | ✓ |
| Year dummies | ✓ | ✓ | ✓ | ✓ |
| Country time trends | ✓ | ✓ | ✓ | ✓ |
| Observations | 1,080 | 1,080 | 540 | 540 |
| R-squared | 0.501 | 0.546 | 0.681 | 0.755 |
| Long-run elasticity | | 1.229*** | | 1.477*** |
| | | (0.256) | | (0.148) |
| CI test 5% critical value | | -3.96 | | -4.15 |

Notes: Estimation of (2) and (3) with OLS and also including country-specific linear time trends. Parentheses include robust standard errors, brackets t-statitics. 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: 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.9: 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 | 36.383 | 0.634 | 44.082 | 0.303 | 277.207 | 0.000 |
| Inverse normal Z | 1.737 | 0.959 | -0.190 | 0.425 | -13.205 | 0.000 |
| | *Variable: openness* | | | | | |
| Inverse chi-squared(40) P | 54.949 | 0.058 | 118.244 | 0.000 | | |
| Inverse normal Z | -1.831 | 0.034 | -6.843 | 0.000 | | |
| | *Variable: union* | | | | | |
| Inverse chi-squared(40) P | 46.757 | 0.215 | 102.239 | 0.000 | | |
| Inverse normal Z | 0.927 | 0.823 | -5.555 | 0.000 | | |
| | *Variable: govsize* | | | | | |
| Inverse chi-squared(40) P | 50.956 | 0.115 | 124.287 | 0.000 | | |
| Inverse normal Z | -1.290 | 0.099 | -6.566 | 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.

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

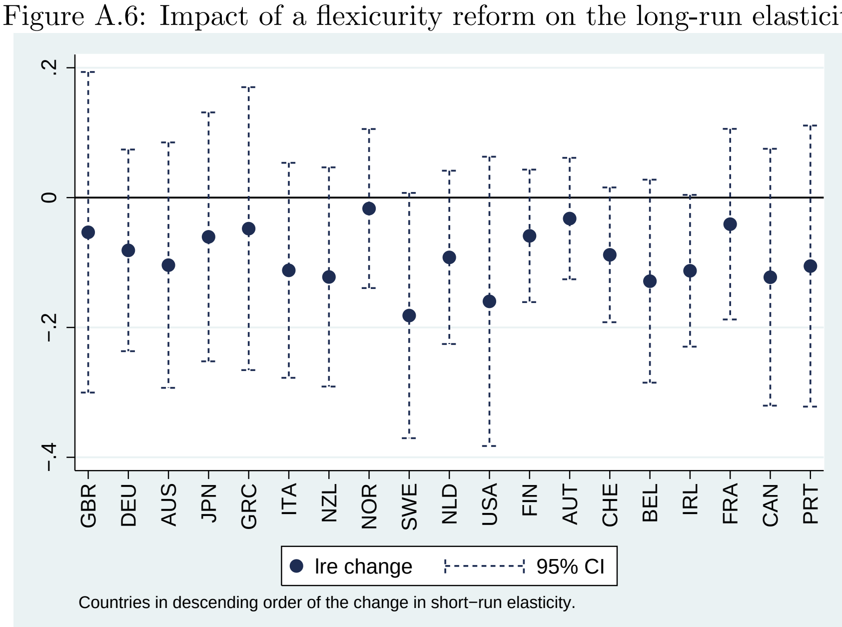

Table A.10: Estimates with policies (EC specification)

| Depvar: $\Delta e$ | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| $\Delta y$ | 0.510*** | 0.531*** | 0.523*** | 0.520*** | 0.451*** | 0.556*** | 0.487*** |
| | (0.042) | (0.047) | (0.045) | (0.042) | (0.046) | (0.038) | (0.042) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ | 0.480* | | | 0.852*** | 0.696** | 0.701*** | 0.619** |
| | (0.270) | | | (0.322) | (0.318) | (0.269) | (0.290) |
| $\Delta y$ x $\tilde{\text{train}}_{-1}$ | | 0.017 | | -0.032 | 0.012 | 0.003 | 0.034 |
| | | (0.027) | | (0.032) | (0.032) | (0.030) | (0.029) |
| $\Delta y$ x $\tilde{\text{flex}}_{-1}$ | | | 0.076** | 0.081** | 0.083* | 0.079*** | 0.085** |
| | | | (0.034) | (0.033) | (0.043) | (0.029) | (0.038) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | 0.299 | | 0.263 |
| | | | | | (0.194) | | (0.175) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.783** | | -0.812** |
| | | | | | (0.376) | | (0.332) |
| $\Delta y$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 0.053 | | 0.029 |
| | | | | | (0.043) | | (0.037) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.244 | | -0.429 |
| | | | | | (0.311) | | (0.269) |
| $e_{-1}$ | -0.070*** | -0.071*** | -0.075*** | -0.084*** | -0.100*** | -0.073*** | -0.077*** |
| | (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 | -0.040 | 0.045 |
| | (0.030) | | | (0.032) | (0.058) | (0.030) | (0.054) |
| $e_{-1}$ x $\tilde{\text{train}}_{-1}$ | | 0.000 | | 0.001 | 0.002 | 0.004 | 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 | | 0.008 |
| | | | | | (0.048) | | (0.045) |
| $e_{-1}$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.030 | | -0.017 |
| | | | | | (0.062) | | (0.061) |
| $e_{-1}$ 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*** | | 0.092** |
| | | | | | (0.049) | | (0.046) |
| $y_{-1}$ | 0.055*** | 0.052*** | 0.052*** | 0.060*** | 0.074*** | 0.047*** | 0.054*** |
| | (0.012) | (0.014) | (0.013) | (0.012) | (0.014) | (0.011) | (0.013) |
| $y_{-1}$ x $\tilde{\text{grr}}_{-1}$ | 0.042 | | | 0.038 | -0.028 | 0.059** | -0.012 |
| | (0.028) | | | (0.030) | (0.050) | (0.028) | (0.046) |
| $y_{-1}$ x $\tilde{\text{train}}_{-1}$ | | 0.000 | | -0.001 | -0.005 | -0.004 | -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 | -0.004 | 0.006 |
| | | | (0.004) | (0.004) | (0.007) | (0.004) | (0.006) |

*Table continues on the next page.*

| Depvar: $\Delta e$ | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| *Continued from the previous page.* | | | | | | | |
| $y_{-1}$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | -0.067 | | -0.036 |
| | | | | | (0.049) | | (0.046) |
| $y_{-1}$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 0.007 | | 0.003 |
| | | | | | (0.057) | | (0.055) |
| $y_{-1}$ 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*** | | -0.067* |
| | | | | | (0.042) | | (0.039) |
| $\tilde{\text{grr}}_{-1}$ | -0.552* | | | -0.514 | 0.193 | -0.736** | 0.026 |
| | (0.316) | | | (0.333) | (0.550) | (0.311) | (0.505) |
| $\tilde{\text{train}}_{-1}$ | | -0.001 | | 0.019 | 0.065 | 0.049 | 0.045 |
| | | (0.027) | | (0.034) | (0.061) | (0.032) | (0.055) |
| $\tilde{\text{flex}}_{-1}$ | | | -0.015 | -0.002 | -0.118 | 0.024 | -0.073 |
| | | | (0.047) | (0.049) | (0.073) | (0.045) | (0.070) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | 0.803 | | 0.438 |
| | | | | | (0.551) | | (0.518) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.046 | | -0.013 |
| | | | | | (0.631) | | (0.609) |
| $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 0.057 | | 0.042 |
| | | | | | (0.049) | | (0.046) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 1.175** | | 0.704 |
| | | | | | (0.461) | | (0.429) |
| $\Delta^2$ wapop | | | | | | 0.357*** | 0.319** |
| | | | | | | (0.124) | (0.125) |
| $\Delta$ openness | | | | | | -0.055 | -0.042 |
| | | | | | | (0.035) | (0.033) |
| $\Delta$ union | | | | | | -0.256*** | -0.259*** |
| | | | | | | (0.054) | (0.057) |
| $\Delta$ govsize | | | | | | 0.548*** | 0.496*** |
| | | | | | | (0.098) | (0.101) |
| Country dummies | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Year dummies | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Observations | 516 | 516 | 516 | 516 | 516 | 516 | 516 |
| R-squared | 0.680 | 0.670 | 0.676 | 0.697 | 0.732 | 0.744 | 0.770 |
| Wald test on joint significance of policy interactions: | | | | | | | |
| F-statistic | | | | | 2.79 | | 2.23 |
| p-value | | | | | 0.0003 | | 0.0041 |

Notes: OLS estimation of (4) 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. *** p<0.01, ** p<0.05, * p<0.1.

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

| Depvar: $\Delta e$ | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| $\Delta y$ | 0.523*** | 0.532*** | 0.527*** | 0.530*** | 0.483*** | 0.552*** | 0.507*** |
| | (0.046) | (0.048) | (0.047) | (0.045) | (0.047) | (0.040) | (0.040) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ | 0.502* | | | 0.869** | 0.760** | 0.641** | 0.652** |
| | (0.283) | | | (0.340) | (0.345) | (0.290) | (0.308) |
| $\Delta y$ x $\tilde{\text{train}}_{-1}$ | | 0.022 | | -0.036 | 0.011 | -0.000 | 0.029 |
| | | (0.028) | | (0.034) | (0.034) | (0.032) | (0.030) |
| $\Delta y$ x $\tilde{\text{flex}}_{-1}$ | | | 0.068** | 0.080** | 0.055 | 0.083*** | 0.059* |
| | | | (0.033) | (0.033) | (0.040) | (0.030) | (0.035) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | 0.270 | | 0.209 |
| | | | | | (0.201) | | (0.179) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.956*** | | -0.957*** |
| | | | | | (0.369) | | (0.318) |
| $\Delta y$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 0.090** | | 0.055* |
| | | | | | (0.036) | | (0.031) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.026 | | -0.289 |
| | | | | | (0.292) | | (0.255) |
| $\tilde{\text{grr}}_{-1}$ | -0.027** | | | -0.037** | -0.038** | -0.023* | -0.028** |
| | (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 | -0.002 |
| | | | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | -0.012 | | -0.013 |
| | | | | | (0.009) | | (0.008) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.012 | | -0.013 |
| | | | | | (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** | | -0.023* |
| | | | | | (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 | 0.719 | 0.741 |
| Wald test on joint significance of policy interactions: | | | | | | | |
| F-statistic | | | | | 3.48 | | 4.13 |
| p-value | | | | | 0.0007 | | 0.0001 |

Notes: Estimation of (4) 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. Control variables include the annual changes in trade openness, trade union density, government size and the change in the working-age population growth rate. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

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

| Depvar: $\Delta e$ | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| $\Delta y$ | 0.500*** | 0.509*** | 0.497*** | 0.509*** | 0.455*** | 0.543*** | 0.488*** |
| | (0.043) | (0.044) | (0.043) | (0.043) | (0.043) | (0.039) | (0.038) |
| $\Delta y$ x $\tilde{grr}_{-1}$ | 0.518** | | | 0.575* | 0.573* | 0.408 | 0.454 |
| | (0.263) | | | (0.301) | (0.310) | (0.260) | (0.278) |
| $\Delta y$ x $\tilde{train}_{-1}$ | | 0.043 | | 0.005 | 0.036 | 0.035 | 0.056** |
| | | (0.028) | | (0.036) | (0.031) | (0.032) | (0.028) |
| $\Delta y$ x $\tilde{flex}_{-1}$ | | | 0.092*** | 0.091*** | 0.081** | 0.082*** | 0.081** |
| | | | (0.033) | (0.032) | (0.039) | (0.029) | (0.034) |
| $\Delta y$ x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ | | | | | 0.308 | | 0.244 |
| | | | | | (0.197) | | (0.172) |
| $\Delta y$ x $\tilde{grr}_{-1}$ x $\tilde{flex}_{-1}$ | | | | | -0.552 | | -0.623** |
| | | | | | (0.357) | | (0.302) |
| $\Delta y$ x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$ | | | | | 0.052 | | 0.025 |
| | | | | | (0.039) | | (0.034) |
| $\Delta y$ x $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$ | | | | | -0.223 | | -0.449* |
| | | | | | (0.310) | | (0.269) |
| $\tilde{grr}_{-1}$ | -0.038** | | | -0.037** | -0.048** | -0.024 | -0.031* |
| | (0.017) | | | (0.016) | (0.020) | (0.014) | (0.017) |
| $\tilde{train}_{-1}$ | | 0.001 | | 0.002 | 0.001 | 0.001 | -0.001 |
| | | (0.001) | | (0.002) | (0.001) | (0.001) | (0.001) |
| $\tilde{flex}_{-1}$ | | | -0.005 | -0.008** | -0.002 | -0.006* | -0.001 |
| | | | (0.003) | (0.004) | (0.004) | (0.003) | (0.004) |
| $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ | | | | | -0.004 | | -0.008 |
| | | | | | (0.010) | | (0.009) |
| $\tilde{grr}_{-1}$ x $\tilde{flex}_{-1}$ | | | | | -0.033* | | -0.027* |
| | | | | | (0.018) | | (0.016) |
| $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$ | | | | | -0.005** | | -0.004** |
| | | | | | (0.002) | | (0.002) |
| $\tilde{grr}_{-1}$ x $\tilde{train}_{-1}$ x $\tilde{flex}_{-1}$ | | | | | -0.017 | | -0.012 |
| | | | | | (0.013) | | (0.013) |
| Control variables | | | | | | ✓ | ✓ |
| Country dummies | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Year dummies | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Country-spec. trends | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Observations | 516 | 516 | 516 | 516 | 516 | 516 | 516 |
| R-squared | 0.699 | 0.698 | 0.702 | 0.713 | 0.732 | 0.757 | 0.776 |
| Wald test on joint significance of policy interactions: | | | | | | | |
| F-statistic | | | | | 3.01 | | 3.82 |
| p-value | | | | | 0.0026 | | 0.0002 |

Notes: OLS estimation of (4) with country-specific linear time trends also included. 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. Control variables include the annual changes in trade openness, trade union density, government size and the change in the working-age population growth rate. *** p<0.01, ** p<0.05, * p<0.1.

Table A.13: Effects on the long-run elasticity of $\sigma$-increases in the policy variables

| Policy variable | without interactions | with interactions average | flexible | generous |
|---|---|---|---|---|
| grr | 0.054*** | 0.036 | 0.045 | 0.009 |
| | (0.019) | (0.030) | (0.031) | (0.032) |
| train | -0.024 | -0.045 | 0.020 | -0.073* |
| | (0.019) | (0.028) | (0.067) | (0.041) |
| flex | 0.032 | 0.030 | 0.057 | 0.029 |
| | (0.025) | (0.025) | (0.044) | (0.038) |

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.10. Average, flexible and generous 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.14: Robustness: Effects on the short-run elasticity of $\sigma$-increases in the policy variables

| Policy variable | without interactions | with interactions average | flexible | generous |
|---|---|---|---|---|
| *First-difference specification* | | | | |
| grr | 0.080** | 0.081** | -0.009 | 0.200*** |
| | (0.036) | (0.038) | (0.053) | (0.063) |
| train | 0.000 | 0.033 | 0.087 | 0.026 |
| | (0.036) | (0.033) | (0.126) | (0.040) |
| flex | 0.082*** | 0.058* | 0.103*** | 0.006 |
| | (0.029) | (0.034) | (0.037) | (0.051) |
| *First-difference specification with country trends* | | | | |
| grr | 0.051 | 0.057 | 0.004 | 0.160*** |
| | (0.032) | (0.035) | (0.051) | (0.057) |
| train | 0.040 | 0.063** | 0.143 | 0.093** |
| | (0.036) | (0.032) | (0.125) | (0.043) |
| flex | 0.081*** | 0.080** | 0.102*** | 0.024 |
| | (0.029) | (0.034) | (0.034) | (0.050) |

Notes: The figures show the effects of a 1-standard-deviation increase in a policy variable on the short-run employment-output elasticity. In the first panel, estimates without interactions are based on column (6), those with interactions on column (7) of Table A.11. In the second panel, estimates are based on the same columns of Table A.12. Average, flexible and generous 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.15: Policy distance from Denmark in 2010

| Country | Policy variable | | |
|---|---|---|---|
| code (ISO) | 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.

Table A.16: Impact of flexicurity reform: comparing different estimates

| Country code | EC (baseline) | | | FD | | | FD with country trends | | |
|---|---|---|---|---|---|---|---|---|---|
| | sre change | s.e. | | sre change | s.e. | | sre change | s.e. | |
| GBR | 0.260 | 0.086 | *** | 0.273 | 0.082 | *** | 0.330 | 0.083 | *** |
| DEU | 0.240 | 0.069 | *** | 0.227 | 0.073 | *** | 0.222 | 0.069 | *** |
| AUS | 0.217 | 0.074 | *** | 0.218 | 0.074 | *** | 0.255 | 0.073 | *** |
| JPN | 0.207 | 0.066 | *** | 0.205 | 0.067 | *** | 0.245 | 0.065 | *** |
| GRC | 0.193 | 0.129 | | 0.170 | 0.132 | | 0.222 | 0.122 | * |
| ITA | 0.179 | 0.061 | *** | 0.141 | 0.061 | ** | 0.203 | 0.062 | *** |
| NZL | 0.164 | 0.065 | * | 0.160 | 0.065 | ** | 0.180 | 0.064 | *** |
| NOR | 0.153 | 0.066 | ** | 0.159 | 0.066 | ** | 0.137 | 0.063 | ** |
| SWE | 0.152 | 0.080 | * | 0.124 | 0.082 | | 0.190 | 0.080 | ** |
| NLD | 0.145 | 0.048 | *** | 0.128 | 0.049 | *** | 0.154 | 0.049 | *** |
| USA | 0.140 | 0.131 | | 0.194 | 0.118 | | 0.184 | 0.121 | |
| FIN | 0.115 | 0.037 | *** | 0.107 | 0.039 | *** | 0.108 | 0.037 | *** |
| AUT | 0.112 | 0.043 | *** | 0.110 | 0.045 | ** | 0.089 | 0.041 | ** |
| CHE | 0.084 | 0.039 | ** | 0.075 | 0.040 | * | 0.081 | 0.039 | ** |
| BEL | 0.084 | 0.058 | | 0.051 | 0.059 | | 0.121 | 0.058 | ** |
| IRL | 0.080 | 0.069 | | 0.082 | 0.068 | | 0.089 | 0.065 | |
| FRA | 0.059 | 0.075 | | 0.047 | 0.075 | | 0.078 | 0.074 | |
| CAN | 0.052 | 0.080 | | 0.046 | 0.081 | | 0.082 | 0.076 | |
| PRT | 0.026 | 0.094 | | -0.042 | 0.090 | | 0.086 | 0.090 | |

Notes: Estimated changes in the short-run employment-output elasticity and the corresponding standard errors for three different model specification. The calculations are based on column 7 of Tables A.10, A.11 and A.12, respectively. Countries are listed from the highest to the lowest change according to the baseline estimates. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

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

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**Abstract**

Labour market reforms in the direction of 'flexicurity' have been widely endorsed as a means to increasing 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 panel of OECD countries, which also incorporates the period of the Great Recession, and distinguish between passive and active labour market policy types. We find that the effects of any single policy change are shaped by the broader existing policy-mix within which it takes place.

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

**JEL codes:** E24, E32, J21, J65

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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 these inter-country variations 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. 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. 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.[2] 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 *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

[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.*, 2020), 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.*, 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 *et al.*, 2014).

been forcefully demonstrated theoretically – e.g. Davoine and Keuschnigg (2015), Dabusinskas *et al.* (2016) and Molana *et al.* (2020) – 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.

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. The long-run elasticity, however, has 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 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*, *flexible* and *generous*. 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.* (2018, 2020), 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 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., 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) 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 cycles, 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 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) 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

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 (e.g., Hendry, 1995). An advantage of this 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.[3] 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 *et al.* (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 ($y$) 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

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

$t$ 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 bracket) and a random error term ($\epsilon_{i,t}$) capturing contemporaneous employment shocks. The term in the bracket is the long-run equilibrium relationship between employment and output, where $\beta$ is assumed to be negative and $-\beta$ measures the long-run employment to output elasticity. The long-run relationship also allows for a deterministic linear time trend, which may represent long-run technological or demographic changes that are common across countries. The short-run employment-output elasticity is captured by $\gamma$.

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 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 $\alpha$.

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.

We reformulate equation (1) into the following estimating 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 $\alpha$ remains the speed of adjustment parameter, $\gamma$ is the short-run elasticity, while the long-run elasticity is obtained as $-\frac{\beta^*}{\alpha}$. Estimating this equation instead of (1) allows for more flexibility in the treatment of the common time trend, as it assumes year-specific fixed effects ($\delta_t$) instead of a linear time trend. Moreover, the equation accounts for country-specific constants ($\delta_i$). As a robustness check, we will also experiment with adding country-specific linear time trends, which can account for the possibility that secular trends, e.g. in terms of demography or technology, differ from country to country.

In the above framework, it is possible to assess the existence of a cointegrating relationship by testing whether $\alpha$ is significantly smaller than zero (Ericsson and MacKinnon, 2002). Should the estimated value of $\alpha$ be non-negative, then the cointegrating relationship is considered as nonexistent, 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 $\gamma$. 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 *et al.*, 2012; Bassanini, 2012) and include interactions of the right-hand side variables with the policy variables in the estimating 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 the ease of interpretation, the policy variables are 'centered' within each year, i.e. expressed as deviations from their country means as $\tilde{P}_{i,t} = P_{i,t} - \frac{1}{N}\sum_{j=1}^{N} P_{j,t}$. 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,

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

where $k$ is the index for the different policy measures and – possibly – their interactions. The estimate for $\gamma_1$ will always give the short-run elasticity for the average country.[4] The estimated $\gamma_2$ 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,

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

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

Policy variables can, of course, relate endogeneously 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. 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., union density, openness to international trade or the size of the government, which 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 the employment responsiveness (e.g., Duval *et al.*, 2020; Ciminelli *et al.*, 2018). 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* instead 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 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 employ-

[5]The Penn World Table we use is described in Feenstra *et al.* (2015). The data is available for download at `www.ggdc.net/pwt`.

ment and output. 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 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 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 differences.

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 exists, which we test in two ways. 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 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 examinig the statistical significance of the speed of adjustment parameter when estimating equation (2) 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.[6]

## 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 (2) and (3) 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 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 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 estimate of the long-run employment-output elasticity, obtained using error correction specification, is around 0.8 for the average country on both samples. This means that if output permanently increases by 1%, employment will also rise by 0.8% in the long run, ceteris paribus. 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 instantaneous 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.[7]

[6]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.

[7]The response function on Figure A.1 is produced by making dynamic forecasts of our estimated model with and

Table 1: Elasticity estimates

| Depvar: $\Delta e$ | 1960-2014 | | 1986-2012 | |
|---|---|---|---|---|
| $\Delta y$ (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 models (2) and (3) are based on OLS. 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.

Our elasticity estimates are roughly consistent with those found for developed countries in previous literature (e.g. Kapsos, 2006; Crivelli *et al.*, 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.

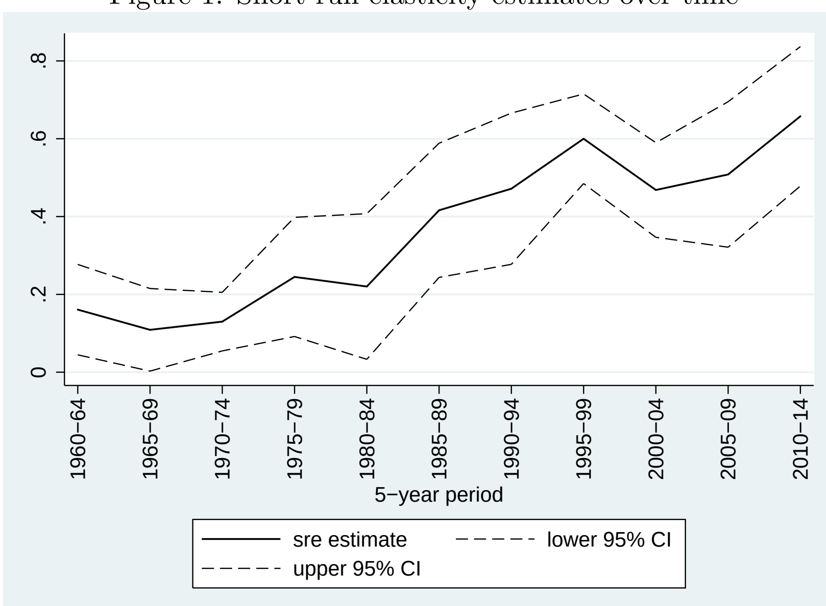
A robustness exercise is reported in Table A.7 which shows that the short-run elasticity estimates are remarkably robust to the inclusion of country-specific linear time trends. We also find evidence that the cointegrating relationship exists under the assumption of such trends and that - at least on the long panel - the speed of adjustment parameter and the long-run elasticity do not change significantly compared to the baseline estimates. Nevertheless, on the short panel, the latter two estimates 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 specific trends in the error correction (EC) regression on the short panel carries the risk of model overfitting. For this reason, in what follows we only

without a 1% shock to output and taking the difference of the two forecast paths.

include country specific trends in the more parsimonious first differenced (FD) specification.

Next, we investigate more closely how the short-run elasticity changes over time. To this end, we interact the output change variable with (five-year) period dummies when estimating equation (2) on the long panel. The resulting period-specific elasticity estimates are plotted in Figure 1.[8] 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 – for instance, as suggested by Blanchard *et al.* (2006) it 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.

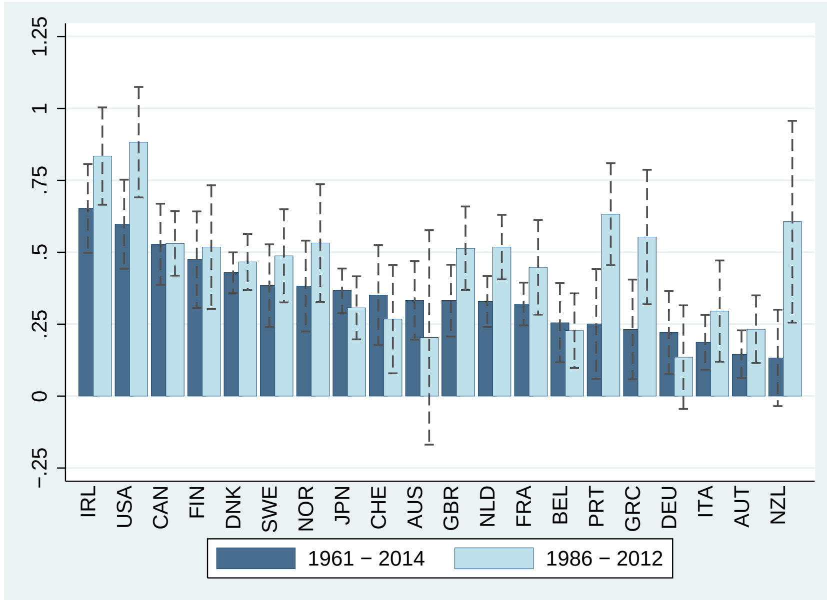
Figure 1: Short-run elasticity estimates over time




In order to establish the extent of inter-country heterogeneity in employment responsiveness, we estimate model (2) by interacting the output change variable with country dummies to obtain country-specific elasticities. Figure 2 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. It is apparent that considerable inter-country variations exist in the responsiveness of employment to output. Countries with the largest elastic-

[8]Estimating the FD specification (3) or including country-specific trends in the regression result in very similar graphs.

Figure 2: Estimated short-run elasticities by country




ities – such as the US and Ireland on the left-hand side of the graph – have at least three times higher employment responsiveness than countries on the right-hand side (e.g. Austria, Germany). The estimates for some countries also change considerably when the sample is shortened to the more recent decades. In particular, the elasticities for Greece, New Zealand and Portugal, which all have below-average elasticities in the long panel, increase to above average.[9]

In the remaining part of this paper, we focus on the role of labour market institutions and policies in driving inter-country heterogenous responses in employment-output elasticities.

## 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 considered as 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 needs for the flexibility required to adjust to negative shocks with that for adequate

[9]These findings are largely unchanged (though with some changes in the country ranking) when estimating the first-difference specification or including country-specific trends in the regression.

social safety nets (see, e.g., Blanchard *et al.*, 2014). The descriptive statistics of the three policy variables are reported in Table 2 below.

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.[10] *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.[11]

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.2). 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 in both hiring and firing is empirically challenging. However, arguably – as is well demonstrated in the theoretical literature (see, for

[10] 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.

[11] 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.

instance, Zanetti, 2011; Cacciatore and Fiori, 2016) – the 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.[12] Our *flex* index measures, on a 0 to 6 scale, how easy it is for employers to lay off workers (firing) or to employ them on temporary work contracts (hiring), as laid down by a country's regulatory system.[13] Hence, a higher value of *flex* means a more flexible labour market, which entail both lower costs of hiring and firing. 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 *targeted* 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 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

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

[13] 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* variable.

suggests the relatively 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.8 reveal that countries with generous unemployment benefits tend to 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 (Blanchard *et al.*, 2014; 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).

Thus, we estimate the EC equation while introducing the policy variables and their interactions with the right-hand-side variables and report the estimation results in Table A.10. The coefficient estimates in the different columns of the table are reasonably stable, regardless whether policies are included individually or jointly and whether policy interactions are also accounted for. This stability also holds for the inclusion of country-time-specific control variables, which is shown in the last two columns. The set of control variables includes the size of the working-age population, openness to international trade, trade union density, and the size of the government (see a more detailed description in Table A.1), that can potentially affect employment elasticities. Panel unit root tests indicate that all of these variables are first-order integrated except for the working-age population, which is second-order integrated (Table A.9). This motivates us to include in the regression the annual changes of trade openness, trade union density and government size and the change in the annual change (second-order difference) of the working-age population. All but trade openness turn out to be significant predictors of employment growth. At the same time, their inclusion leaves the rest of the coefficient estimates roughly unchanged.

The estimates in Table A.10 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 figure of 0.701 in the second row of column 6, 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 0.07 higher-than-average short-run elasticity. Similarly, a one-point higher flexibility index corresponds to a 0.08 higher elasticity. 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.

Columns 5 and 7 consider all possible interactions between the policy variables. 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 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.

The above results are largely unchanged when we estimate the FD specification (Table A.11) or the FD specification with country-specific linear time trends (Table A.12). A minor difference is that in the FD regression the interaction of *train* and *flex* is significantly positively associated with the short-run employment to output elasticity. This finding, however, disappears when country-specific time trends are also included.

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 *et al.*, 2018; Shimer, 2005; Zanetti, 2011) as our estimates reveal. 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 job 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 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 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 greater 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). The offsetting effects of unemployment generosity and flexibility produces 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 (e.g. Onwordi, 2016; Pissarides, 2009; Stähler and Thomas, 2012), a positive shock to productivity is likely to result in greater elasticity in the short-run since it reduces the profitability of job matches. However, as argued by Molana *et al.* (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 supports 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*, *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 (5) to calculate the effects on the short-run elasticity, using the estimated coefficients in Table A.10, reported either in column 6 (without policy interactions) or in column 7 (with policy interactions). 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 $\hat{\gamma}_2^{\tilde{\mathrm{grr}}} \times \hat{\sigma}^{\mathrm{grr}} = 0.701 \times 0.125 = 0.088$, where $\hat{\gamma}_2^{\tilde{\mathrm{grr}}}$ is the estimated coefficient for the interaction of $\Delta y$ with the gross replacement rate in column 6, and $\hat{\sigma}^{\mathrm{grr}}$ 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

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

where the $\hat{\gamma}_2$s are the coefficients for the corresponding policy interactions with $\Delta y$ in column 7 of Table A.10 and the policy variables $\tilde{\mathrm{train}}$ and $\tilde{\mathrm{flex}}$ 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.

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 0.088 and 0.078, respectively). 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 *flexible* regimes, making the unemployment benefit more

Table 3: Effects on the short-run elasticity of $\sigma$-increases in the policy variables

| Policy variable | without interactions | with interactions average | flexible | generous |
|---|---|---|---|---|
| grr | 0.088*** | 0.077** | 0.004 | 0.197*** |
| | (0.034) | (0.036) | (0.055) | (0.058) |
| train | 0.003 | 0.039 | 0.102 | 0.066 |
| | (0.033) | (0.033) | (0.137) | (0.043) |
| flex | 0.078*** | 0.083** | 0.122*** | 0.017 |
| | (0.029) | (0.037) | (0.038) | (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. Estimates without interactions are based on column (6), those with interactions on column (7) of Table A.10. *Average*, *flexible* and *generous* 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.

generous does not result in 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 a *generous* 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 *et al.* (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. As evident in column 2 (where complementarities are not taken into account), however, implementing each of these policies in an average regime, will have significant effects. 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 robust to changing the model specification into FD with or without country-specific trends (Table A.14).

Finally, we calculate the corresponding changes in the long-run elasticity, which we report in Table A.13.[14] The numbers suggest that the effects of these policies found on the short-run elasticity

[14] This we do by evaluating expression (6) before and after the one-stdev 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.

mostly fade out in the long run. Training is the only policy with a (weakly) significant and negative long-run impact on the employment elasticities 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.*, 2018).[15]

## 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-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.15 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 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, 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 3, in descending order, together with the 95% confidence intervals. The computation was based on expression (5) and the estimated coefficients in Table A.10, column 7,

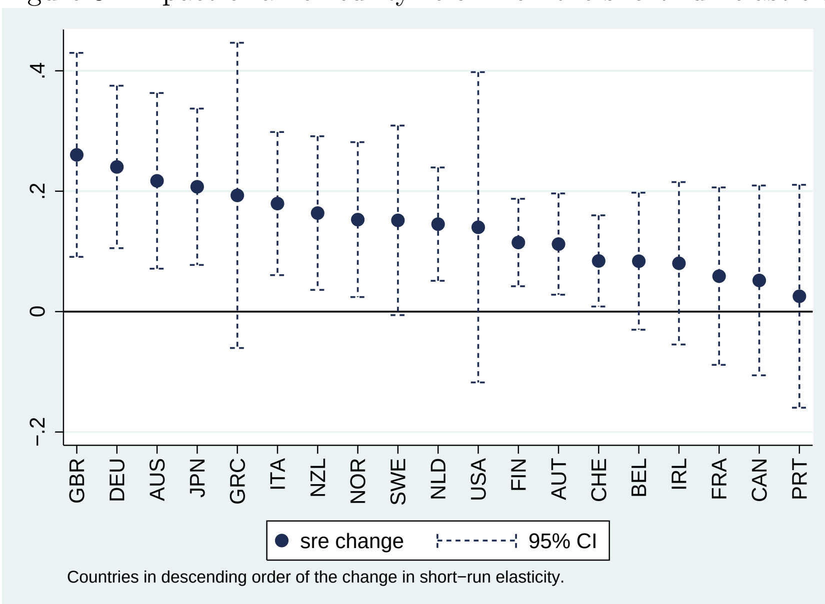
[15] In a search and matching open economy model, Molana *et al.* (2018) show that increases in training expenditure can reduce employment volatility in response to exogenous shocks even in the presence of more generous unemployment insurance.

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 significant effects are obtained also for Italy, New Zealand, Norway, the Netherlands, Finland, Austria and Switzerland. At the other end, the estimated effects are small or not statistically significant for Portugal, Canada, France, Ireland, Belgium, the US and Greece.

The group of countries with the highest expected effects includes those with 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 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, either with or without country trends. Table A.16 compares the baseline estimates with the results of the two robustness checks. The groups of countries for which statistically significant increases in the short-run elasticity are predicted are almost identical under the three specifications, and the relative sizes of the country estimates also remain largely unchanged.

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 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

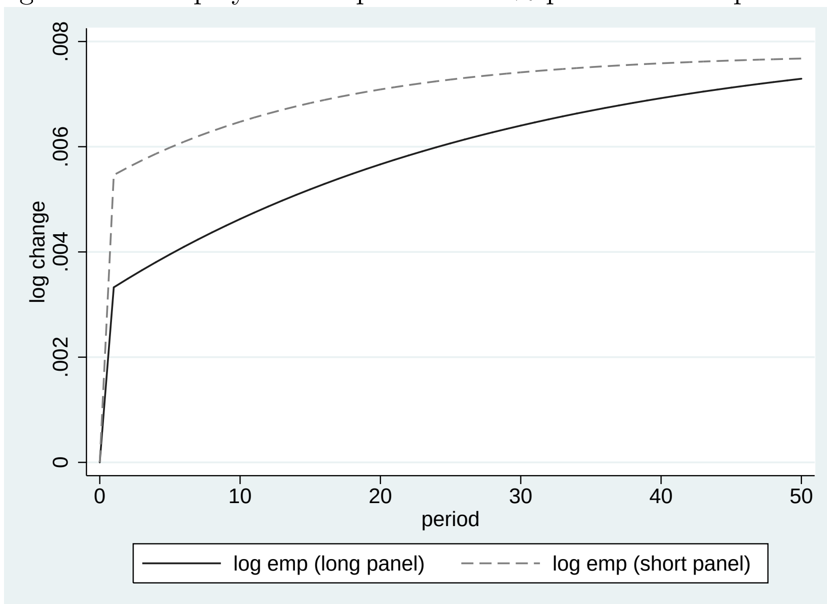


Figure A.2: Benefit replacement rate by country

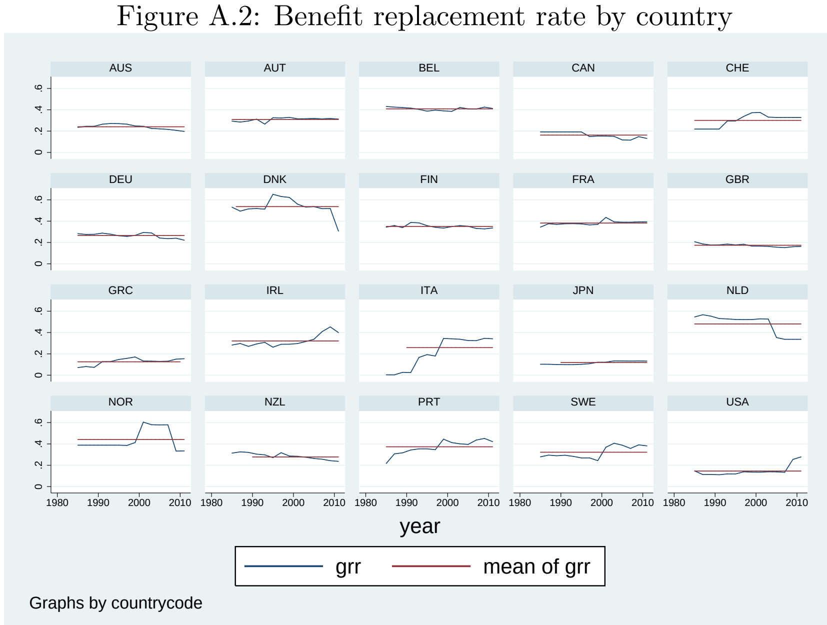

Table A.1: Variable description

| Variable | Description | Data Source |
|---|---|---|
| | *Core variables* | |
| e (employment) | Number of persons engaged (millions, in log) | Penn World Table 9.0 |
| y (output) | Real GDP at constant national prices (mil. 2011US$, in log) | Penn World Table 9.0 |
| | *Policy variables* | |
| grr | Gross unemployment benefit replacement rate | OECD |
| train | Training expenditures per unemployed person (PPP US$, log) | OECD |
| flex | Labour market flexibility indicator (0-6 scale) | OECD, EPL indicators |
| | *Control variables* | |
| wapop | Working-age population (ages 20-64, mil. persons, in log) | OECD |
| openness | Trade openness (export+import)/GDP in real 2010US$ | World Bank, WDI |
| union | Trade union density (share) | OECD |
| govsize | Government consumption as share of GDP | OECD, National Accounts |

Notes: The Penn World Table 9.0 is described in Feenstra *et al.* (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 | -1.444 | 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 ($\Delta e$) | | | | | dlog output ($\Delta y$) | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| 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.

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

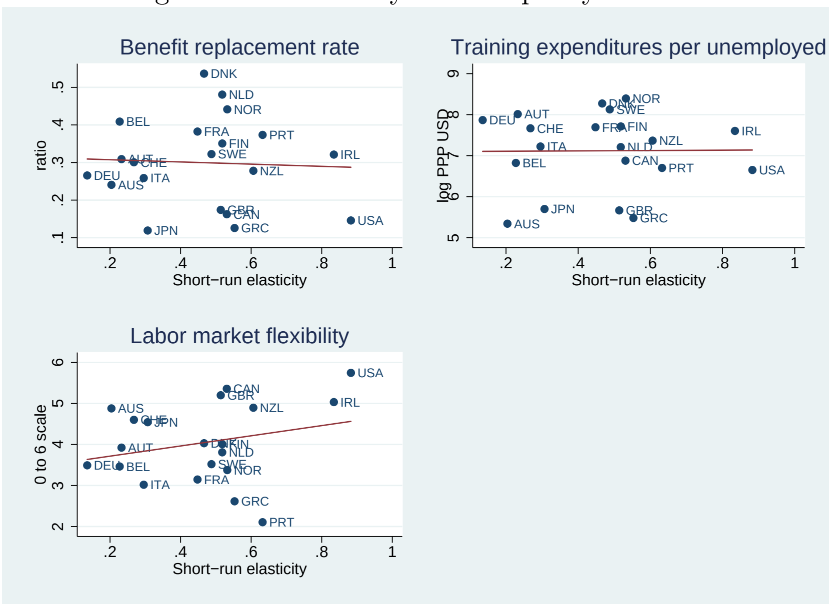


Table A.5: Panel unit root tests for first differences

| | dlog employment ($\Delta e$) | | dlog output ($\Delta y$) | |
|---|---|---|---|---|
| | 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.

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: Robustness: Elasticity estimates with country trends

| Depvar: $\Delta e$ | 1960-2014 | | 1986-2012 | |
|---|---|---|---|---|
| $\Delta y$ (Short-run elasticity) | 0.325*** | 0.345*** | 0.509*** | 0.535*** |
| | (0.034) | (0.032) | (0.040) | (0.035) |
| $e_{-1}$ (Speed of adjustment) | | -0.047*** | | -0.151*** |
| | | (0.011) | | (0.022) |
| | | [-4.24] | | [-6.76] |
| $y_{-1}$ | | 0.058*** | | 0.222*** |
| | | (0.008) | | (0.022) |
| Country dummies | ✓ | ✓ | ✓ | ✓ |
| Year dummies | ✓ | ✓ | ✓ | ✓ |
| Country time trends | ✓ | ✓ | ✓ | ✓ |
| Observations | 1,080 | 1,080 | 540 | 540 |
| R-squared | 0.501 | 0.546 | 0.681 | 0.755 |
| Long-run elasticity | | 1.229*** | | 1.477*** |
| | | (0.256) | | (0.148) |
| CI test 5% critical value | | -3.96 | | -4.15 |

Notes: Estimation of (2) and (3) with OLS and also including country-specific linear time trends. Parentheses include robust standard errors, brackets t-statitics. 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: 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.9: 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 | 36.383 | 0.634 | 44.082 | 0.303 | 277.207 | 0.000 |
| Inverse normal Z | 1.737 | 0.959 | -0.190 | 0.425 | -13.205 | 0.000 |
| *Variable: openness* | | | | | | |
| Inverse chi-squared(40) P | 54.949 | 0.058 | 118.244 | 0.000 | | |
| Inverse normal Z | -1.831 | 0.034 | -6.843 | 0.000 | | |
| *Variable: union* | | | | | | |
| Inverse chi-squared(40) P | 46.757 | 0.215 | 102.239 | 0.000 | | |
| Inverse normal Z | 0.927 | 0.823 | -5.555 | 0.000 | | |
| *Variable: govsize* | | | | | | |
| Inverse chi-squared(40) P | 50.956 | 0.115 | 124.287 | 0.000 | | |
| Inverse normal Z | -1.290 | 0.099 | -6.566 | 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.

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

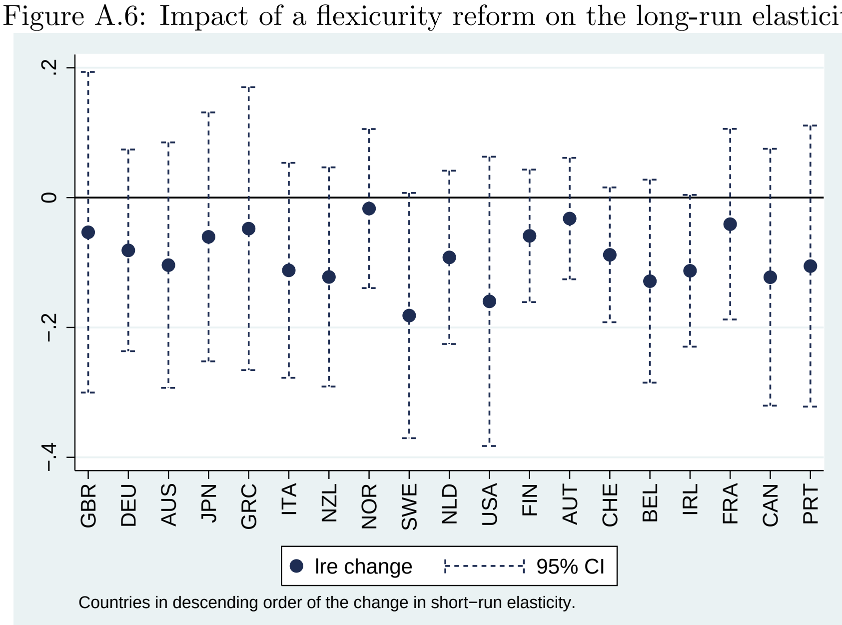

Table A.10: Estimates with policies (EC specification)

| Depvar: $\Delta e$ | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| $\Delta y$ | 0.510*** | 0.531*** | 0.523*** | 0.520*** | 0.451*** | 0.556*** | 0.487*** |
| | (0.042) | (0.047) | (0.045) | (0.042) | (0.046) | (0.038) | (0.042) |
| $\Delta y$ x $\widetilde{\text{grr}}_{-1}$ | 0.480* | | | 0.852*** | 0.696** | 0.701*** | 0.619** |
| | (0.270) | | | (0.322) | (0.318) | (0.269) | (0.290) |
| $\Delta y$ x $\widetilde{\text{train}}_{-1}$ | | 0.017 | | -0.032 | 0.012 | 0.003 | 0.034 |
| | | (0.027) | | (0.032) | (0.032) | (0.030) | (0.029) |
| $\Delta y$ x $\widetilde{\text{flex}}_{-1}$ | | | 0.076** | 0.081** | 0.083* | 0.079*** | 0.085** |
| | | | (0.034) | (0.033) | (0.043) | (0.029) | (0.038) |
| $\Delta y$ x $\widetilde{\text{grr}}_{-1}$ x $\widetilde{\text{train}}_{-1}$ | | | | | 0.299 | | 0.263 |
| | | | | | (0.194) | | (0.175) |
| $\Delta y$ x $\widetilde{\text{grr}}_{-1}$ x $\widetilde{\text{flex}}_{-1}$ | | | | | -0.783** | | -0.812** |
| | | | | | (0.376) | | (0.332) |
| $\Delta y$ x $\widetilde{\text{train}}_{-1}$ x $\widetilde{\text{flex}}_{-1}$ | | | | | 0.053 | | 0.029 |
| | | | | | (0.043) | | (0.037) |
| $\Delta y$ x $\widetilde{\text{grr}}_{-1}$ x $\widetilde{\text{train}}_{-1}$ x $\widetilde{\text{flex}}_{-1}$ | | | | | -0.244 | | -0.429 |
| | | | | | (0.311) | | (0.269) |
| $e_{-1}$ | -0.070*** | -0.071*** | -0.075*** | -0.084*** | -0.100*** | -0.073*** | -0.077*** |
| | (0.013) | (0.014) | (0.014) | (0.015) | (0.016) | (0.014) | (0.015) |
| $e_{-1}$ x $\widetilde{\text{grr}}_{-1}$ | -0.020 | | | -0.017 | 0.064 | -0.040 | 0.045 |
| | (0.030) | | | (0.032) | (0.058) | (0.030) | (0.054) |
| $e_{-1}$ x $\widetilde{\text{train}}_{-1}$ | | 0.000 | | 0.001 | 0.002 | 0.004 | 0.001 |
| | | (0.003) | | (0.003) | (0.006) | (0.003) | (0.005) |
| $e_{-1}$ x $\widetilde{\text{flex}}_{-1}$ | | | 0.003 | 0.004 | -0.011 | 0.009* | -0.005 |
| | | | (0.005) | (0.005) | (0.008) | (0.005) | (0.007) |
| $e_{-1}$ x $\widetilde{\text{grr}}_{-1}$ x $\widetilde{\text{train}}_{-1}$ | | | | | 0.033 | | 0.008 |
| | | | | | (0.048) | | (0.045) |
| $e_{-1}$ x $\widetilde{\text{grr}}_{-1}$ x $\widetilde{\text{flex}}_{-1}$ | | | | | -0.030 | | -0.017 |
| | | | | | (0.062) | | (0.061) |
| $e_{-1}$ x $\widetilde{\text{train}}_{-1}$ x $\widetilde{\text{flex}}_{-1}$ | | | | | 0.009* | | 0.008* |
| | | | | | (0.005) | | (0.004) |
| $e_{-1}$ x $\widetilde{\text{grr}}_{-1}$ x $\widetilde{\text{train}}_{-1}$ x $\widetilde{\text{flex}}_{-1}$ | | | | | 0.136*** | | 0.092** |
| | | | | | (0.049) | | (0.046) |
| $y_{-1}$ | 0.055*** | 0.052*** | 0.052*** | 0.060*** | 0.074*** | 0.047*** | 0.054*** |
| | (0.012) | (0.014) | (0.013) | (0.012) | (0.014) | (0.011) | (0.013) |
| $y_{-1}$ x $\widetilde{\text{grr}}_{-1}$ | 0.042 | | | 0.038 | -0.028 | 0.059** | -0.012 |
| | (0.028) | | | (0.030) | (0.050) | (0.028) | (0.046) |
| $y_{-1}$ x $\widetilde{\text{train}}_{-1}$ | | 0.000 | | -0.001 | -0.005 | -0.004 | -0.004 |
| | | (0.002) | | (0.003) | (0.005) | (0.003) | (0.005) |
| $y_{-1}$ x $\widetilde{\text{flex}}_{-1}$ | | | 0.000 | -0.001 | 0.010 | -0.004 | 0.006 |
| | | | (0.004) | (0.004) | (0.007) | (0.004) | (0.006) |

*Table continues on the next page.*

| Depvar: $\Delta e$ | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| *Continued from the previous page.* | | | | | | | |
| $y_{-1}$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | -0.067 | | -0.036 |
| | | | | | (0.049) | | (0.046) |
| $y_{-1}$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 0.007 | | 0.003 |
| | | | | | (0.057) | | (0.055) |
| $y_{-1}$ 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*** | | -0.067* |
| | | | | | (0.042) | | (0.039) |
| $\tilde{\text{grr}}_{-1}$ | -0.552* | | | -0.514 | 0.193 | -0.736** | 0.026 |
| | (0.316) | | | (0.333) | (0.550) | (0.311) | (0.505) |
| $\tilde{\text{train}}_{-1}$ | | -0.001 | | 0.019 | 0.065 | 0.049 | 0.045 |
| | | (0.027) | | (0.034) | (0.061) | (0.032) | (0.055) |
| $\tilde{\text{flex}}_{-1}$ | | | -0.015 | -0.002 | -0.118 | 0.024 | -0.073 |
| | | | (0.047) | (0.049) | (0.073) | (0.045) | (0.070) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | 0.803 | | 0.438 |
| | | | | | (0.551) | | (0.518) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.046 | | -0.013 |
| | | | | | (0.631) | | (0.609) |
| $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 0.057 | | 0.042 |
| | | | | | (0.049) | | (0.046) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 1.175** | | 0.704 |
| | | | | | (0.461) | | (0.429) |
| $\Delta^2$ wapop | | | | | | 0.357*** | 0.319** |
| | | | | | | (0.124) | (0.125) |
| $\Delta$ openness | | | | | | -0.055 | -0.042 |
| | | | | | | (0.035) | (0.033) |
| $\Delta$ union | | | | | | -0.256*** | -0.259*** |
| | | | | | | (0.054) | (0.057) |
| $\Delta$ govsize | | | | | | 0.548*** | 0.496*** |
| | | | | | | (0.098) | (0.101) |
| Country dummies | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Year dummies | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Observations | 516 | 516 | 516 | 516 | 516 | 516 | 516 |
| R-squared | 0.680 | 0.670 | 0.676 | 0.697 | 0.732 | 0.744 | 0.770 |
| Wald test on joint significance of policy interactions: | | | | | | | |
| F-statistic | | | | | 2.79 | | 2.23 |
| p-value | | | | | 0.0003 | | 0.0041 |

Notes: OLS estimation of (4) 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. *** p<0.01, ** p<0.05, * p<0.1.

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

| Depvar: $\Delta e$ | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| $\Delta y$ | 0.523*** | 0.532*** | 0.527*** | 0.530*** | 0.483*** | 0.552*** | 0.507*** |
| | (0.046) | (0.048) | (0.047) | (0.045) | (0.047) | (0.040) | (0.040) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ | 0.502* | | | 0.869** | 0.760** | 0.641** | 0.652** |
| | (0.283) | | | (0.340) | (0.345) | (0.290) | (0.308) |
| $\Delta y$ x $\tilde{\text{train}}_{-1}$ | | 0.022 | | -0.036 | 0.011 | -0.000 | 0.029 |
| | | (0.028) | | (0.034) | (0.034) | (0.032) | (0.030) |
| $\Delta y$ x $\tilde{\text{flex}}_{-1}$ | | | 0.068** | 0.080** | 0.055 | 0.083*** | 0.059* |
| | | | (0.033) | (0.033) | (0.040) | (0.030) | (0.035) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | 0.270 | | 0.209 |
| | | | | | (0.201) | | (0.179) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.956*** | | -0.957*** |
| | | | | | (0.369) | | (0.318) |
| $\Delta y$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 0.090** | | 0.055* |
| | | | | | (0.036) | | (0.031) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.026 | | -0.289 |
| | | | | | (0.292) | | (0.255) |
| $\tilde{\text{grr}}_{-1}$ | -0.027** | | | -0.037** | -0.038** | -0.023* | -0.028** |
| | (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 | -0.002 |
| | | | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | -0.012 | | -0.013 |
| | | | | | (0.009) | | (0.008) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.012 | | -0.013 |
| | | | | | (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** | | -0.023* |
| | | | | | (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 | 0.719 | 0.741 |
| Wald test on joint significance of policy interactions: | | | | | | | |
| F-statistic | | | | | 3.48 | | 4.13 |
| p-value | | | | | 0.0007 | | 0.0001 |

Notes: Estimation of (4) 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. Control variables include the annual changes in trade openness, trade union density, government size and the change in the working-age population growth rate. *** p<0.01, ** p<0.05, * p<0.1.

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

| Depvar: $\Delta e$ | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| $\Delta y$ | 0.500*** | 0.509*** | 0.497*** | 0.509*** | 0.455*** | 0.543*** | 0.488*** |
| | (0.043) | (0.044) | (0.043) | (0.043) | (0.043) | (0.039) | (0.038) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ | 0.518** | | | 0.575* | 0.573* | 0.408 | 0.454 |
| | (0.263) | | | (0.301) | (0.310) | (0.260) | (0.278) |
| $\Delta y$ x $\tilde{\text{train}}_{-1}$ | | 0.043 | | 0.005 | 0.036 | 0.035 | 0.056** |
| | | (0.028) | | (0.036) | (0.031) | (0.032) | (0.028) |
| $\Delta y$ x $\tilde{\text{flex}}_{-1}$ | | | 0.092*** | 0.091*** | 0.081** | 0.082*** | 0.081** |
| | | | (0.033) | (0.032) | (0.039) | (0.029) | (0.034) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | 0.308 | | 0.244 |
| | | | | | (0.197) | | (0.172) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.552 | | -0.623** |
| | | | | | (0.357) | | (0.302) |
| $\Delta y$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | 0.052 | | 0.025 |
| | | | | | (0.039) | | (0.034) |
| $\Delta y$ x $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.223 | | -0.449* |
| | | | | | (0.310) | | (0.269) |
| $\tilde{\text{grr}}_{-1}$ | -0.038** | | | -0.037** | -0.048** | -0.024 | -0.031* |
| | (0.017) | | | (0.016) | (0.020) | (0.014) | (0.017) |
| $\tilde{\text{train}}_{-1}$ | | 0.001 | | 0.002 | 0.001 | 0.001 | -0.001 |
| | | (0.001) | | (0.002) | (0.001) | (0.001) | (0.001) |
| $\tilde{\text{flex}}_{-1}$ | | | -0.005 | -0.008** | -0.002 | -0.006* | -0.001 |
| | | | (0.003) | (0.004) | (0.004) | (0.003) | (0.004) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ | | | | | -0.004 | | -0.008 |
| | | | | | (0.010) | | (0.009) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.033* | | -0.027* |
| | | | | | (0.018) | | (0.016) |
| $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.005** | | -0.004** |
| | | | | | (0.002) | | (0.002) |
| $\tilde{\text{grr}}_{-1}$ x $\tilde{\text{train}}_{-1}$ x $\tilde{\text{flex}}_{-1}$ | | | | | -0.017 | | -0.012 |
| | | | | | (0.013) | | (0.013) |
| | | | | | | | |
| Control variables | | | | | | ✓ | ✓ |
| Country dummies | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Year dummies | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Country-spec. trends | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| | | | | | | | |
| Observations | 516 | 516 | 516 | 516 | 516 | 516 | 516 |
| R-squared | 0.699 | 0.698 | 0.702 | 0.713 | 0.732 | 0.757 | 0.776 |
| Wald test on joint significance of policy interactions: | | | | | | | |
| F-statistic | | | | | 3.01 | | 3.82 |
| p-value | | | | | 0.0026 | | 0.0002 |

Notes: OLS estimation of (4) with country-specific linear time trends also included. 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. Control variables include the annual changes in trade openness, trade union density, government size and the change in the working-age population growth rate. *** p<0.01, ** p<0.05, * p<0.1.

Table A.13: Effects on the long-run elasticity of $\sigma$-increases in the policy variables

| Policy variable | without interactions | with interactions average | flexible | generous |
|---|---|---|---|---|
| grr | 0.054*** | 0.036 | 0.045 | 0.009 |
| | (0.019) | (0.030) | (0.031) | (0.032) |
| train | -0.024 | -0.045 | 0.020 | -0.073* |
| | (0.019) | (0.028) | (0.067) | (0.041) |
| flex | 0.032 | 0.030 | 0.057 | 0.029 |
| | (0.025) | (0.025) | (0.044) | (0.038) |

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.10. Average, flexible and generous 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.14: Robustness: Effects on the short-run elasticity of $\sigma$-increases in the policy variables

| Policy variable | without interactions | with interactions average | flexible | generous |
|---|---|---|---|---|
| *First-difference specification* | | | | |
| grr | 0.080** | 0.081** | -0.009 | 0.200*** |
| | (0.036) | (0.038) | (0.053) | (0.063) |
| train | 0.000 | 0.033 | 0.087 | 0.026 |
| | (0.036) | (0.033) | (0.126) | (0.040) |
| flex | 0.082*** | 0.058* | 0.103*** | 0.006 |
| | (0.029) | (0.034) | (0.037) | (0.051) |
| *First-difference specification with country trends* | | | | |
| grr | 0.051 | 0.057 | 0.004 | 0.160*** |
| | (0.032) | (0.035) | (0.051) | (0.057) |
| train | 0.040 | 0.063** | 0.143 | 0.093** |
| | (0.036) | (0.032) | (0.125) | (0.043) |
| flex | 0.081*** | 0.080** | 0.102*** | 0.024 |
| | (0.029) | (0.034) | (0.034) | (0.050) |

Notes: The figures show the effects of a 1-standard-deviation increase in a policy variable on the short-run employment-output elasticity. In the first panel, estimates without interactions are based on column (6), those with interactions on column (7) of Table A.11. In the second panel, estimates are based on the same columns of Table A.12. Average, flexible and generous 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.15: Policy distance from Denmark in 2010

| Country | Policy variable | | |
|---|---|---|---|
| code (ISO) | 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.

Table A.16: Impact of flexicurity reform: comparing different estimates

| Country code | EC (baseline) | | | FD | | | FD with country trends | | |
|---|---|---|---|---|---|---|---|---|---|
| | sre change | s.e. | | sre change | s.e. | | sre change | s.e. | |
| GBR | 0.260 | 0.086 | *** | 0.273 | 0.082 | *** | 0.330 | 0.083 | *** |
| DEU | 0.240 | 0.069 | *** | 0.227 | 0.073 | *** | 0.222 | 0.069 | *** |
| AUS | 0.217 | 0.074 | *** | 0.218 | 0.074 | *** | 0.255 | 0.073 | *** |
| JPN | 0.207 | 0.066 | *** | 0.205 | 0.067 | *** | 0.245 | 0.065 | *** |
| GRC | 0.193 | 0.129 | | 0.170 | 0.132 | | 0.222 | 0.122 | * |
| ITA | 0.179 | 0.061 | *** | 0.141 | 0.061 | ** | 0.203 | 0.062 | *** |
| NZL | 0.164 | 0.065 | * | 0.160 | 0.065 | ** | 0.180 | 0.064 | *** |
| NOR | 0.153 | 0.066 | ** | 0.159 | 0.066 | ** | 0.137 | 0.063 | ** |
| SWE | 0.152 | 0.080 | * | 0.124 | 0.082 | | 0.190 | 0.080 | ** |
| NLD | 0.145 | 0.048 | *** | 0.128 | 0.049 | *** | 0.154 | 0.049 | *** |
| USA | 0.140 | 0.131 | | 0.194 | 0.118 | | 0.184 | 0.121 | |
| FIN | 0.115 | 0.037 | *** | 0.107 | 0.039 | *** | 0.108 | 0.037 | *** |
| AUT | 0.112 | 0.043 | *** | 0.110 | 0.045 | ** | 0.089 | 0.041 | ** |
| CHE | 0.084 | 0.039 | ** | 0.075 | 0.040 | * | 0.081 | 0.039 | ** |
| BEL | 0.084 | 0.058 | | 0.051 | 0.059 | | 0.121 | 0.058 | ** |
| IRL | 0.080 | 0.069 | | 0.082 | 0.068 | | 0.089 | 0.065 | |
| FRA | 0.059 | 0.075 | | 0.047 | 0.075 | | 0.078 | 0.074 | |
| CAN | 0.052 | 0.080 | | 0.046 | 0.081 | | 0.082 | 0.076 | |
| PRT | 0.026 | 0.094 | | -0.042 | 0.090 | | 0.086 | 0.090 | |

Notes: Estimated changes in the short-run employment-output elasticity and the corresponding standard errors for three different model specification. The calculations are based on column 7 of Tables A.10, A.11 and A.12, respectively. Countries are listed from the highest to the lowest change according to the baseline estimates. *** p<0.01, ** p<0.05, * p<0.1.