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Youth Unemployment and Stigmatization Over the Business Cycle in Europe*

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Abstract

This paper studies the dynamics of the scarring effect of youth unemployment over the business cycle in 12 European countries. On the one hand, we analyse differences associated with the negative effect of past unemployment experiences on future labour market status. And, on the other hand, we consider the potential stigmatization of prospective young workers – that is, the extent to which employers are more reluctant to hire individuals with a history of unemployment. Our results are based on data from the EU-SILC for the period 2004 to 2015 and provide support in favour of a significant scarring effect of unemployment among youths that is highly heterogeneous across the countries under analysis and that increased substantially during the Great Recession. In contrast, the evidence of stigma effects was found to be rather weak.

I. Introduction

This paper studies the dynamics of the scarring effect of unemployment over the business cycle. In particular, we analyse differences associated with the negative effect of past unemployment experiences on future labour market status among relevant European Union (EU) countries from 2004 to 2015, a timespan that includes periods of low and relatively stable unemployment rates, as well as the Great Recession. There is a vast corpus of literature showing that unemployment suffers from a considerable degree of persistence (Narendranathan and Elias, 1993; Arulampalam *et al.*, 2000, 2001; Stewart, 2007). On the one hand, unemployment persistence can be explained by

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observed and unobserved characteristics which persist across time and that make someone more likely to be unemployed successively. For example, poor qualifications, low motivation or a general lack of ability (Biewen and Steffes, 2010). On the other hand, it has also been shown that a spell of unemployment increases by itself the likelihood that someone will suffer unemployment again in the future; this has come to be known among labour economists as 'true' or 'genuine' state dependence in unemployment. The sources or mechanisms behind the scarring effect of unemployment have proved to be more difficult to disentangle, but several suggestions have been made: loss of human capital (Pissarides, 1992), unemployment insurance disincentives (Gangl, 2006), minimum wages (Cockx and Ghirelli, 2016), decline in search intensity (Vishwanath, 1989; Cockx and Dejemeppe, 2012), habituation (Clark *et al*., 2001), discouragement (Ayllón, 2013), rational herding[1] (Manning, 2000; Oberholzer-Gee, 2008) and also stigmatization by employers (Lockwood, 1991; Omori, 1997; Biewen and Steffes, 2010).

Stigmatization in the labour market refers to a certain type of discrimination against prospective workers that occurs when employers are more reluctant to hire individuals with a history of long-term unemployment or who have had frequent periods of unemployment.[2] Instead, employers favour individuals who move from job to job or who have had short (and infrequent) spells of unemployment. A theoretical explanation for firms' hiring decisions is that labour markets are characterized by incomplete information, and thus employers rely on signals to estimate the applicants' expected productivity (Hujer *et al*., 2004). Stigmatization may occur because employers regard unemployment as a negative signal, leading them to believe that the individual-specific level of human capital deteriorates while a person is jobless;[3] or they may simply assume that those unemployed individuals are less motivated or less productive (Blau and Robins, 1990; Lockwood, 1991; Omori, 1997; Clark *et al*., 2001).

Previous literature has pointed out that the stigmatization of unemployed individuals is particularly prevalent during periods of economic growth (when the unemployment rate is low). Biewen and Steffes (2010) show that when the unemployment rate rises, state dependence in unemployment decreases, indicating that employers are less suspicious of unemployed individuals during periods of economic downturn, when the unemployment rate is above its trend. By contrast, they discriminate against individuals who are unemployed when the current unemployment rate is low – so in periods of economic growth. In other words, employers become more suspicious of individuals who are unemployed when the economy is growing and unemployment is less prevalent. From the perspective of a hiring firm, the negative signal of unemployment might be especially relevant in the case of prospective young workers, since, compared

[1]Rational herding refers to the idea that managers believe that unemployed applicants must have been previously interviewed, and if the applicants were productive, they would already have been employed (Oberholzer-Gee, 2008).

[2]Another potential consequence of individuals having previous spells of unemployment could be a wage penalty when they are hired (particularly in a growing economy and among individuals with less human capital and/or less experience). Such analysis, although, is beyond the scope of our paper.

[3]Acemoglu (1995) argues that persistence in unemployment is a result of workers not being able to demonstrate convincingly that they have managed to maintain their skills during unemployment.

to their working colleagues of the same age, they lack the positive signal of employment and, moreover, their level of human capital is lower than that of prime-aged job seekers; this signal intensifies in periods of economic upswing, when young workers should normally transit easily into employment (Axelrad *et al*., 2018).

However, research on the analysis of stigmatization in the labour market has mainly focused on prime-aged men, and so we know very little about the (potential) stigmatization of young unemployed individuals or women.[4] Moreover the high levels of joblessness among young people brought about by the Great Recession in Europe also justify analysis of the state dependence and stigmatization of this age group, even if transitions into and out of employment are much more frequent and quick, and sometimes occur for reasons other than the labour market – for example, willingness to pursue further education. The same is true of women, but with additional considerations, as they work part-time more often and their employment histories are more likely to be interrupted by periods of maternal leave or voluntary inactivity to engage in non-market household production. Therefore, employers might consider that these interruptions could have a negative effect on their relative levels of productivity. Also, there has been little or no analysis of the effect of different institutional and national contexts on the stigmatization (if any) of prospective young workers in the labour market. Moreover, as far as we know, stigmatization was barely analysed during the period of the Great Recession – the only exception being Tumino (2015), who shows for the United Kingdom that state dependence in unemployment increases when the rate of unemployment rises. Thus, to the best of our knowledge, this is the first study using a relevant group of EU countries to analyse potential stigma effects among young people from a cross-country perspective, and while considering the extent to which such disadvantage can also affect women.

Our main findings, based on data for 12 European countries and using dynamic random-effects probit models, provide evidence of a significant scarring effect of unemployment among young people. The effect is highly heterogeneous across the European countries under analysis and increased substantially during the Great Recession when measured by average partial effects (APEs). Comparing the results with those for prime-aged workers, our findings suggest that young individuals suffer from higher unemployment rates, but are less affected by state dependence in unemployment in most countries. This reflects the fact that young people enter and leave the labour market more quickly and more often, and are therefore less likely to accumulate long periods of unemployment. As for the existence of potential stigma effects, we only find such a counter-cyclical relationship in Belgium – one of the most rigid labour markets in the OECD (Cockx and Ghirelli, 2016). On the other hand, when we restrict our analysis to the period from 2008 to 2015, the results show that the number of countries with such counter-cyclical association increased – although effects are weakly significant in most contexts.

[4]Previous literature has focused on prime-aged men, because they form the population group that is expected to have the strongest attachment to the labour market, and so an understanding of transitions into and out of employment is a relevant research question with important policy implications.

The paper is organized as follows. Section II provides a literature review. Section III describes our data. Section IV presents the methodology used. Section V shows the results, as well as some robustness checks. And finally, the last section contains a conclusion and a discussion of avenues for future research.

II. Literature review

There is an on-going debate in the literature concerning the extent to which unemployment persistence is due to genuine state dependence or is related to unobserved heterogeneity (see e.g. the literature review in Krueger *et al*., 2014). In the main, there is no consensus on the relative importance of these two factors.

One strand of the literature relies on conditioning on observables to account for differences between those unemployed and those employed at time *t*. Krueger *et al*. (2014) and Kroft *et al*. (2016) use calibrated matching models for the US economy, which account for changes in the composition of the unemployed, based on observed characteristics. They show that workers with an extended period of unemployment have poorer prospects of finding a job, because their skills erode while they are unemployed. The main limitation of these studies is that they cannot account for unobserved heterogeneity, which might affect the relative employability of those unemployed individuals seeking a job on different labour markets. Depending on (un) observed heterogeneity among workers, individuals with long-term histories of unemployment may well face more difficulties in a tight labour market (they might be essentially unemployable), than in the aftermath of a recession.

There is another strand of the literature that uses resume audit studies to analyse the effect of unemployment histories on employer call-back rates. By means of randomized field experiments, these analyses are able to account for individual unobserved heterogeneity. However, they have also yielded mixed results. Kroft *et al*. (2013) and Ghayad (2014), focusing on younger workers with different levels of education in the United States, find that the longer the spell of unemployment, the greater the (negative) effect on the probability of a job applicant receiving a call-back: this is most likely to come in the first year of unemployment. Furthermore, Kroft *et al*. (2013) show that the local labour market conditions (measured in metropolitan statistical areas) have a significant impact on the duration dependence, indicating that tighter labour markets reduce the probability of individuals exiting unemployment. By contrast, also in the United States, Farber *et al*. (2016, 2017), focusing on older college graduates, and Nunley *et al*. (2017), who concentrate on younger college graduates, find no effect of unemployment duration on call-back rates. As for evidence in Europe, Eriksson and Rooth (2014), show that in the case of highly educated Swedish workers, a spell of unemployment has no impact on employers' hiring decisions, whereas it does have a negative effect for less-educated workers. Nüß (2018) finds similar results for younger secondary-school graduates in Germany. Finally, Farber *et al*. (2019), aiming to reconcile the contrasting results on unemployment duration found in the literature, develop a field experiment that covers a broad range of ages. They determine that both younger and older applicants experience a lower call-back probability than prime-aged applicants, although they also show that there is no relationship between current unemployment histories and call-back rates. They draw special attention to the fact that small differences in labour markets, occupations and resume designs might create substantial variations in call-back rates, imposing limits on the external validity of such studies.

A further strand of literature uses reduced-form econometric models to focus on the potential heterogeneity of the treatment effect of unemployment, caused by genuine state dependence, unobserved heterogeneity and differences in labour market conditions. These studies regard as particularly relevant the potential heterogeneity of state dependence over the business cycle. Using data from the British Household Panel Survey (BHPS) and considering the local unemployment-to-vacancy ratio as an inverse measure of labour market tightness, Arulampalam *et al*. (2000) find that in the UK youth unemployment is independent of the business cycle, whereas for men aged 25 and over the likelihood of unemployment increases with a deterioration in the local labour market conditions.[5] Analysing more recent data for the United Kingdom, Tumino (2015) also looks at the relationship between state dependence in unemployment and the business cycle, controlling for local labour market conditions through a measure of the proportion of claimants of unemployment-related benefits in the population aged 16–64. His study focuses on the role of local labour market conditions on the persistence of unemployment, and finds evidence that unemployment scarring affects both young and older men. In contrast to previous studies, his estimates increase when unemployment rises, and fall when the labour market conditions are more favourable, indicating a positive association between the scarring effect of unemployment and the level of unemployment.[6] In particular, he shows that young people were the worst affected during the Great Recession. Thus, his results show a negative association between the business cycle and state dependence in the early 1990s, early 2000s and during the Great Recession.

Additionally, stigmatization by employers has been analysed in the literature as one of the mechanisms determining the occurrence of genuine state dependence (Lockwood, 1991; Omori, 1997; Biewen and Steffes, 2010). The study of the potential stigmatization of the unemployed further examines genuine state dependence and its relationship to the business cycle. This line of research explores the effect of previous unemployment status as a negative signal that increases the risk of individuals staying or becoming unemployed in future periods. Moreover, depending on the macroeconomic conditions, stigma effects differ in the way they are associated with the dynamics of unemployment scarring. Thus, individuals will become stigmatized if they remain unemployed when the economy is growing and total unemployment is low. This is presumably a consequence of employers discounting these individuals' status in the labour market due to the high degree of uncertainty about their

[5]Note that Arulampalam *et al*. (2000) focus on analysing the individual effect of local labour market tightness on the probability of unemployment, rather than looking at variations of state dependence throughout the business cycle, as we do here.

[6]It is important to note that this result does not come from interpretation of the interaction between state dependence and local labour market conditions, which is actually not statistically significant. Rather the author claims this relationship by evaluating the corresponding average partial effects on lagged unemployment over the distribution of the claimant proportion discretized in 2 percentage point bands.

productivity, thus implying a lower labour market value attached to their characteristics. Omori (1997) reports evidence of stigma effects among young men in the United States, using the local unemployment rate at the time non-employment occurred to proxy the general economic conditions. This implies that, when they make hiring decisions, employers will only consider informative the level of unemployment associated with past unemployment spells. Drawing on data from the National Longitudinal Survey of Labor Market Experience Youth Study (NLSY), Omori (1997) finds that workers who were unemployed when labour market conditions were relatively favourable are more severely stigmatized.

Closely related to our study, Biewen and Steffes (2010) analyse stigmatization effects using data from the German Socio-Economic Panel (SOEP). Following Lockwood (1991), and using a broad measure of unemployment, they find that state dependence in unemployment decreases (increases) when the unemployment rate is high (low).[7] On the other hand, when they test the hypothesis suggested by Omori (1997) and consider the level of past unemployment as a measure of the business cycle, no significant effects are found. Thus, they conclude that there is evidence that the scarring effects of unemployment are counter-cyclical in Germany, and interpret their results as weak confirmation of stigma effects among men.

Ayllón (2013) extends this approach by showing that discouragement among unemployed individuals is not constant over the business cycle, as is assumed in Biewen and Steffes (2010). She argues that discouragement will be associated with a fall in search intensity, especially when employment conditions deteriorate. Drawing on the Spanish component of the European Community Household Panel (ECHP), she reports evidence of stigma effects and discouragement explaining persistence of unemployment, with a consequent positive relationship between the scarring effect and the unemployment cycle. In particular, the effect of discouragement runs counter to the decline in stigmatization effects when workers face more favourable labour market conditions.

In an analysis of state dependence in unemployment and stigmatization over the business cycle, it is also relevant to consider potential variability across regions. In this regard, Lindbeck *et al*. (1999) developed a theoretical model that considers the existence of significant regional divergence in economic development and labour market opportunities within a country. Their results reveal that higher regional unemployment rates (as measured by a high proportion of government transfer recipients) should reduce the detrimental effect of unemployment (consistent with low levels of individual disutility). Hence, unfavourable economic conditions should weaken the social stigma associated with unemployment. Lupi and Ordine (2002) find evidence of this hypothesis when they compare northern and southern regions of Italy. They show that individual unemployment experiences may be less important in determining wages in a social environment that is characterized by high unemployment rates and in the presence of a less developed productive structure. This might have an

[7]Lockwood (1991) developed a matching model, where unemployment duration is a signal of productivity and employers use information regarding the current level of unemployment to infer workers' characteristics.

effect by inducing hysteresis of unemployment and reducing the downward pressure of unemployment on wages at the macro level.

Our study contributes to this literature by combining the most significant characteristics of the aforementioned approaches. On the one hand, we present novel cross-country evidence of the scarring effect of unemployment, controlling for the most important observable characteristics at the corresponding stage of the labour market career. Furthermore, our dynamic random-effects probit models allow us to account for individual specific unobserved heterogeneity, alongside regional fixed effects that capture unobservable time-invariant regional characteristics. On the other hand, we study the effect of unemployment histories at different phases of the business cycle, which allows us both to mitigate an important downside associated with resume audit studies – that is, the hiring decisions of potential employers could well be affected by different labour market conditions (Farber *et al*., 2019) – and to analyse the stigma effects potentially suffered by the unemployed, allowing for variability across regions. Finally, we further examine our results in the context of country-specific institutional settings.

III. Data

We use data from the longitudinal component of the European Union – Statistics on Income and Living Conditions (EU-SILC) which, at the time of writing, runs from 2004 to 2015. The greatest advantage of the EU-SILC is that it provides detailed socio-economic and demographic information on individual and household characteristics. Moreover, data are meant to be comparable across all the participating countries. Possibly the most important disadvantage is that, in the large majority of countries, individuals are followed for only four consecutive waves, which means that in each survey year, 25% of the sample (which constitutes a rotational group) is replaced by new interviewees. This implies that (at most) we will be observing changes in young people’s labour market status over three consecutive waves.[8]

Our analysis covers a selection of 12 countries (and 90 regions). Because our methodology (see below) requires information at the regional level, we could only work with those countries that provide such a variable in the survey.[9] Moreover, not all the countries started their participation in the survey from 2004, and not all the countries provide the regional information for the whole period under analysis – see all these details in Table A.1 in the Appendix. Despite all these limitations, our sample is composed of countries from Continental Europe (Austria and Belgium), the Mediterranean area (Greece, Spain and Italy) and Central and Eastern Europe (Bulgaria, Czech Republic, Poland and Hungary). The results relative to the Great Recession also refer to Nordic Europe (Sweden) and the English-speaking countries

[8]Our pooled data set has been constructed by taking the information from the last file in which a given rotational group appears (Iacovou and Lynn, 2013), which guarantees that the same survey methodology is applied to a given individual over time.

[9]Finland provides information at the regional level; however, there was a change in 2008, which meant that some region codes disappeared, while others were introduced. Thus, the information is not provided longitudinally.

(United Kingdom); thus, our working sample includes examples of different welfare state regimes and geographical areas.

Our main sample contains young people from the age of 17–29 (and from 30 to 44 when we want to compare the results for young people with those of prime-aged individuals). As shown in the first column of Table A.2 in the Appendix, and in relation to our dependent variable, 12% of young people in the sample were unemployed. However, it is important to note that youth unemployment rates were at very different levels and evolved very differently across the period under analysis. Figure A.1 in the Appendix shows the youth unemployment rates (17–29) between 2005 and 2015, drawn from the EU-SILC for the countries under analysis. As can be seen, it is in Greece, Spain and Italy where the youth unemployment rate increased the most, reaching its maximum value around 2013. In the rest of the countries, the trend remains much more stable, although at different levels, with Austria having the lowest values of all those countries analysed in this paper.

Table A.2 also indicates that, on average, the probability of persisting in unemployment if previously unemployed is 53.6% among young people, while the likelihood of unemployment at t if not unemployed at $t-1$ is only 7.0%. Yet the probability of persistence in unemployment varies greatly across the countries under analysis and also in the different periods under analysis. Table 1 shows precisely the probability of being unemployed at time t conditional on being unemployed in the previous period ($t-1$) for three periods of time: 2004–2007, 2008–2011 and 2012–2015. It can be seen that while persistence in unemployment is relatively low in some countries and did not increase much with the outbreak of the Great Recession (see e.g. the case of Austria), in other countries, persistence in unemployment was high at the beginning of the period and increased even further from 2008 to 2011 and from 2012 to 2015 (Bulgaria or Greece). In other countries (e.g. France and Poland), persistence

TABLE 1

Probability of being unemployed at t *conditional on being unemployed at* t − *1, selected countries, Europe, 2004–2015*

Country	*2004–2007*	*2008–2011*	*2012–2015*
Austria	33.7	36.1	40.2
Belgium	50.1	54.9	50.2
Bulgaria	50.7	56.2	73.3
Czech Rep.	43.4	50.0	47.3
Greece	51.3	58.2	76.6
France	50.0	53.8	55.7
Hungary	34.7	43.9	43.5
Italy	62.9	56.2	60.5
Poland	48.5	45.4	55.1
Spain	37.9	54.9	57.6
Sweden	—	33.2	28.8
UK	—	34.5	43.5

Note: The unemployment persistence rate has been computed at the individual level using microdata from the EU-SILC.
Source: Authors' elaboration. EU-SILC, 2005–2015.

in unemployment among young people was high, but did not increase much with the start of the economic crisis.

As for the control variables used in the analysis, and as shown in Table A.2, average age is 23.3 years, 48.8% of the sample are females, 24.5% have not graduated from a high school or a vocational programme, 52.7% hold a secondary education degree and 22.8% are university graduates. As for demographic characteristics, 41.3% of the sample live outside the parental home, 19.6% have a partner and 11.7% cohabit with their own children.[10] As the rest of the table shows, heterogeneity is important across countries, particularly with respect to demographic characteristics.

IV. Empirical strategy

Basic setup

To analyse the scarring effect of unemployment among young adults, we assume that labour market dynamics follow a first-order Markov process and that the past labour market position has a *genuine* impact on the current position. The country-specific dynamic reduced-form model can be written as follows:

$$y_{itrp} = \mathbf{1}\left(\alpha_1 y_{i(t-1)rp} + x'_{itrp}\beta + \mu_r + \mu_p + \tau_{it} > 0\right), \tag{1}$$

where $i = 1, 2, \ldots, N$ are individuals, $r = 1, \ldots, R$ are regions, p refers to the country-specific period under study (see Table A.1 in the Appendix, second column) and $t = 1, 2, 3$ are the individual-specific time points.[11]

The dependent variable (y_{itrp}) is equal to 1 if the individual i is unemployed at time t, and 0 otherwise (e.g. employed, in education, etc.). Following the assumption of a first-order Markov process, y_{itrp} is explained by its lagged outcome, $y_{i(t-1)rp}$. Furthermore, x_{itrp} is a vector of explanatory variables that include time-invariant (gender) and time-variant characteristics (age, age squared, maximum level of education attained, living outside the parental home, having a partner and the number of own children in the household). μ_r are regional fixed effects that capture individual-independent regional characteristics and μ_p are the year fixed effects that capture events common to all regions within one country. Lastly, τ_{it} is an individual- and time-specific shock. Note that for each individual i at every time point t there is no variation across region r or period p; accordingly, for the sake of simplicity, henceforth we do not use the subscripts r and p other than in the notation for the region and period fixed effects.

Importantly, the parameter α_1 measures the extent to which being unemployed in the previous year has an influence on the likelihood of being unemployed again at time t – that is, it captures the degree of *true* or *genuine* state dependence in

[10]Unfortunately, information on the country of origin is not available in the longitudinal component of the EU-SILC.

[11]As mentioned earlier, the EU-SILC is a rotating panel in which an individual is observed for four consecutive waves, which reduces the dynamic sequence to three time points at most.

unemployment. We expect α_1 to be positive and highly significant among the sample of youths under study. However, if unobserved heterogeneity between individuals is persistent over time, not controlling for this aspect results in overstating the effect of α_1 (Heckman and Borjas, 1980; Flaig *et al.*, 1993; Mühleisen and Zimmerman, 1994; Stewart, 2007). To account for it, we follow an approach used intensively in the economic literature (e.g. Arulampalam, 2001; Biewen and Steffes, 2010; Bhuller *et al.*, 2017), and we apply the following decomposition:

$$\tau_{it} = v_i + \varepsilon_{it}, \tag{2}$$

where v_i corresponds to an individual-specific time-constant unobserved effect and ε_{it} is the idiosyncratic error term (Wooldridge, 2005).[12] Furthermore, we assume that both error terms follow a normal distribution, for example, $v_i \sim N(0, \sigma_v^2)$ and $\varepsilon_{it} \sim N(0, \sigma_\varepsilon^2)$, and that ε_{it} is independent and identically distributed (iid). This results in a standard uncorrelated random-effects model, in which the assumption is that the explanatory variables are uncorrelated with the individual-specific time-invariant error term. However, to assume that unobservable differences between individuals are uncorrelated with observable characteristics might appear unrealistic. For example, someone with a higher innate motivation might be more likely to hold a postgraduate degree. To relax this assumption, we follow the suggestion of Mundlak (1978) and Chamberlain (1984) and allow for the following relationship $v_i = \bar{x}_i'\pi + e_i$ with e_i being iid and $e_i \sim N(0, \sigma_e^2)$ and independent of the covariates and the idiosyncratic shock. Inserting (2) into Equation (1) leads to:

$$y_{it} = \mathbf{1}(\alpha_1 y_{i(t-1)} + x_{it}'\beta + \mu_r + \mu_p + \bar{x}_i'\pi + e_i + \varepsilon_{it} > 0). \tag{3}$$

A second aspect that needs to be addressed is the endogeneity of the initial conditions. Up to now, it has been assumed that the unemployment status at $t = 0$ is exogenously given – for example, the labour market position at the first time point is randomly assigned to an individual. However, due to the individual-specific effects being persistent over time, it is likely that the labour market status observed for the first time is itself a result of the past. Thus, not controlling for the endogeneity of the initial conditions can lead to overestimation of state dependence in unemployment (Chay and Hyslop, 2000). There exist different econometric approaches to account for the initial conditions problem (Heckman, 1981a, 1981b; Arulampalam and Stewart, 2009; Skrondal and Rabe-Hesketh, 2014). Wooldridge (2005) established a simple

[12]We do not control for heterogeneity in the unobservables when experiencing unemployment. To tackle that, one can use heterogeneous slope models, in which a second random-effects error term is included to capture heterogeneity in unemployment experiences and which is only estimated if the individual was unemployed at $t-1$. However, our own simulations indicated that for short panels (such as that used in the current paper), the estimator becomes inefficient. Moreover, the estimator does not converge for several countries. Plum and Ayllón (2015) find that not accounting for the heterogeneous effect of past unemployment underestimates the scarring effect of unemployment. Therefore, the extent of state dependence we find in this study must be considered as a conservative lower boundary. An extension of the heterogeneous slope model would be to interact the second random-effects error term with an indicator for the business cycle moment (e.g. the regional unemployment rate). This would enable us to account further for the changing composition of the pool of the unemployed, and to control for aspects that are not captured in the covariates or the random effects. Yet, given the impossibility of estimating such models in short panels, we consider it a future research task.

approach by conditioning the estimation on the first observation of each individual. That is, instead of finding the density of the dependent variable from $t=0, 1, \ldots, T$ conditioned on the explanatory variables, we find the density of the dependent variable from $t=1, \ldots, T$ conditioned on the initial condition and the explanatory variables. Our individual-specific time-invariant error term now takes the following form:

$$v_i = \bar{x}_i'\pi + a_0 + \gamma y_{i0} + \varphi_i, \tag{4}$$

which inserted into equation (3) gives

$$y_{it} = 1\left(\alpha_1 y_{i(t-1)} + x_{it}'\beta + \mu_r + \mu_p + \bar{x}_i'\pi + a_0 + \gamma y_{i0} + \varphi_i + \varepsilon_{it} > 0\right), \tag{5}$$

with $\varphi_i \; iid \sim N(0, \sigma_\varphi^2)$. As φ_i is time-invariant, the composite error term τ_{it} is correlated over time and the correlation between two (different) time points is constant (Arulampalam and Booth, 1998; Arulampalam, 1999) and takes the following equi-correlation structure:[13]

$$\rho = \operatorname{corr}(\tau_{it}, \tau_{is}) = \frac{\sigma_\varphi^2}{\sigma_\varphi^2 + \sigma_\varepsilon^2} \tag{6}$$

with $t \neq s$ and $t, s = 1, 2, 3$.

Accounting for stigmatization

The basic model in equation (5) assumes that the scarring effect of unemployment is constant throughout the business cycle. To relax this assumption, we follow Biewen and Steffes (2010) and Ayllón (2013) and add an interaction term between the individual past unemployment status ($y_{i(t-1)}$) and the measure of the cyclical unemployment risk in period p at the regional level (ue_{rp}). For this, we compute for each region a simple ordinary least squares regression of the unemployment rate (obtained from the Labour Force Survey database) against time and predict the residual.[14] Figure 1 shows the unemployment rates and the fitted values at the country level – being the variability at the regional level used in the paper even larger (not shown). We also include in our regression the cyclical unemployment risk at period p (ue_{rp}), which controls for the simple fact that there are more (fewer) jobs available when the economy is growing (shrinking).

[13]Note that the parameter ρ represents the proportion of the total variance of the error term accounted for by the unobserved individual heterogeneity. Furthermore, under the assumption of $\rho = 0$, we could estimate a pooled dynamic probit model ignoring the error structure and the panel nature of our data. The results of an estimation that ignores the presence of unobserved heterogeneity would clearly overestimate our findings on genuine state dependence. These results are available from the authors upon request.

[14]We have opted to use the data from the European Union – Labour Force Survey because rates are drawn from a larger number of observations and are therefore less subject to measurement error than if they were derived directly from the EU-SILC.

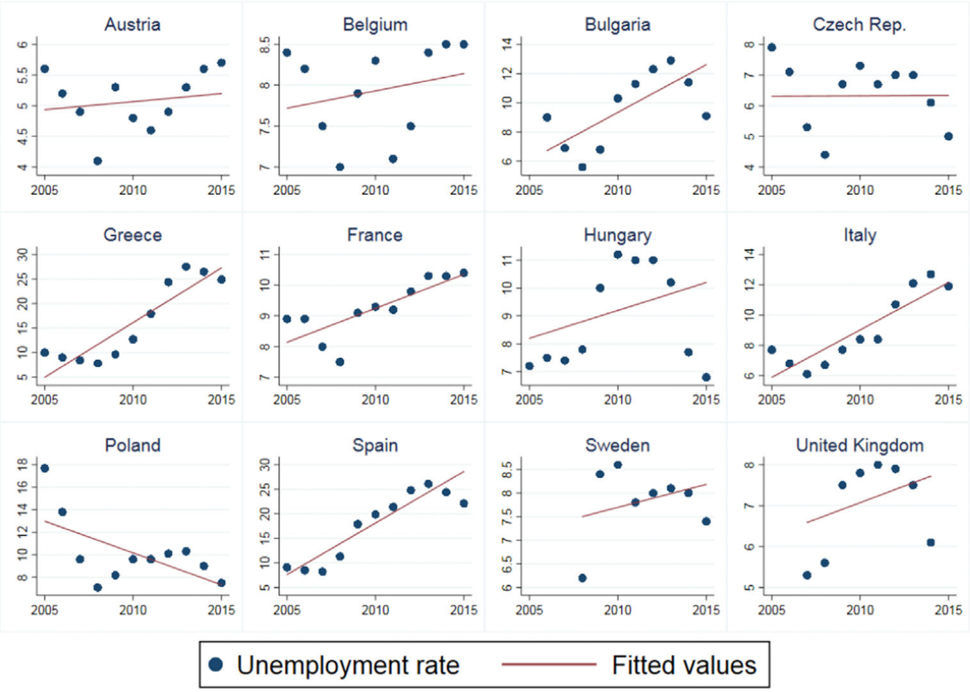

Figure 1. Total unemployment rate and unemployment trend (fitted values), selected countries, Europe, 2005–2015
Source: Authors' elaboration. EU-SILC, 2005–2015.

Formally, our final reduced-form equation can be written as follows:

$$y_{it} = \mathbf{1}\Big(\alpha_1 y_{i(t-1)} + \alpha_2 \mathrm{ue}_{rp} y_{i(t-1)} + \alpha_3 \mathrm{ue}_{rp} + x'_{it}\beta + \mu_r + \mu_p + \bar{x}'_i\pi + a_0 + \gamma y_{i0} + \varphi_i + \varepsilon_{it} > 0\Big), \tag{7}$$

where the individual-specific effect is defined as in equation (4). Our parameters of interest are α_1 for the scarring effect of unemployment (as explained above); α_2 to assess the existence of stigma effects in the labour market; and α_3 for the effect of the business cycle on unemployment.

In particular, if there are stigma effects for prospective workers, we would expect α_2 to be negative, indicating that genuine state dependence decreases when the unemployment rate deviates positively from its trend, which in turn implies that persistence in unemployment is higher when the economy is growing and the unemployment rate is lower – potentially because of stigmatization. In other words, if a counter-cyclical relation is found between state dependence in unemployment and the cyclical unemployment risk ($\alpha_2 < 0$), that would indicate that persistence in unemployment becomes less (more) important in periods of high (low) unemployment; this would indicate that employers are less (more) suspicious of unemployed individuals when unemployment is a more (less) widespread phenomenon. On the other hand, $\alpha_2 > 0$ would indicate that unemployment persistence behaves pro-cyclically: the disadvantage of having been unemployed at $t-1$ is high when the unemployment rate is relatively high, and low when the unemployment rate is relatively low, which could not be attributed to a process of stigmatization. Also, note that α_3 measures the direct impact of the business cycle on the probability of being unemployed, whereas α_2 captures the extent to which the effect of genuine state dependence is heterogeneous throughout the business cycle.

Furthermore, as the outcome variable is dichotomous, a normalization of ε_{it} is required. We assume that $\varepsilon_{it} \sim N(0, 1)$ and the outcome probability is:

$$P_{it}(\varphi^*) = \mathbf{\Phi}\left[\left(\alpha_1 y_{i(t-1)} + \alpha_2 \mathrm{ue}_{rp} y_{i(t-1)} + \alpha_3 \mathrm{ue}_{rp} + x'_{it}\beta + \mu_r + \mu_p + \bar{x}'_i \pi + a_0 + \gamma y_{i0} + \sigma^2_{\varphi}\varphi^*\right)(2y_{it} - 1)\right], \tag{8}$$

where $\mathbf{\Phi}[\bullet]$ refers to the cumulative standard normal distribution, and the likelihood function is the product of all time-point specific probabilities across all individuals. Namely,

$$L = \prod_{i=1}^{N} \int_{\varphi^*} \left\{ \prod_{t=1}^{3} P_{it}(\varphi^*) \right\} \mathrm{d}F(\varphi^*), \tag{9}$$

where F is the distribution function of $\varphi^* = \varphi / \sigma_{\varphi}$. Equation (9) does not have a closed-form solution, and therefore φ has to be integrated out. As we assume that φ is normally distributed, the integral can be evaluated using Gaussian–Hermite quadrature (Butler and Moffitt, 1982).[15]

For the sake of robustness – and following the proposal by Omori (1997) – we have also run specifications that consider the interaction term between past unemployment status of the individual and the measure of the cyclical unemployment rate at $t-1$.[16] That changes equation (7) to

$$y_{it} = \mathbf{1}\left(\alpha_1 y_{i(t-1)} + \alpha_4 \mathrm{ue}_{r(p-1)} y_{i(t-1)} + \alpha_5 \mathrm{ue}_{r(p-1)} + x'_{it}\beta + \mu_r + \mu_p + \bar{x}'_i \pi + a_0 + \gamma y_{i0} + \varphi_i + \varepsilon_{it} > 0\right). \tag{10}$$

The intuition underlying this specification is that what matters to employers is the unemployment risk at the time when past unemployment occurred, rather than the current labour market situation. This means that hiring decisions would be conditioned on the specific circumstances of past unemployment spells: if those were associated with favourable labour market conditions, employers would interpret it as a negative signal, and therefore stigmatization would rise.

[15] We used 12 points, although the main results are not sensitive to change in the number of quadrature points.

[16] Additionally, we performed two other falsification exercises. First, we used the Hodrick–Prescott filter to estimate the regional deviation from the unemployment trend, and our results were almost identical. Second, we randomly swapped the regional unemployment rate within a country. Our results hardly changed. This can be explained by the high correlation of the regional unemployment rate within countries (e.g. corr = 0.79 in our base estimation across all countries).

V. Results

Our findings are presented in six subsections. First, we give details of the main results from the estimation of equation (7). Second, we present results by gender. Third, we engage in a comparative exercise by analysing the situation facing young people in relation to that facing prime-aged individuals. Fourth, we look at whether the situation of young people changed with the outbreak of the Great Recession. Fifth, we present some robustness checks. And, finally, we consider the extent to which the heterogeneity of our results can be related to certain institutional settings.

Main results

Table 2 presents the main results for the countries under analysis. The first column shows the coefficients relative to the parameter α_1 – that is, the degree of genuine state dependence in unemployment. Column (2) presents α_1's associated average partial effect (APE), which is calculated for each individual, holding fixed his or her characteristics, and is averaged over the sample. Column (3) details the results of α_2, which accounts for the interaction between the lagged individual status of unemployment and the cyclical unemployment rate – potentially indicating the existence of stigma effects. And column (4) indicates α_3, which is the estimated coefficient for the cyclical unemployment rate.

In relation to state dependence, the results are very clear. Youth unemployment suffers from an important degree of genuine state dependence, whereby being unemployed at $t-1$ by itself increases the probability of being unemployed again in the following year. However, important differences in terms of the level can be observed across the countries analysed. It is in Bulgaria, Greece and Poland that state dependence is strongest: being unemployed at $t-1$ increases the probability of being unemployed at t by nearly 33 percentage points in Bulgaria. The figures for Greece and Poland are 23 and 21 percentage points respectively. Belgium, France, Italy and Spain lie in the middle of the rankings, with figures of below 20 but above 10 percentage points. And the rest of the countries yield estimates of below 10 percentage points; the figure is particularly low in the Czech Republic (slightly above 5 percentage points). Note that from the results presented we cannot establish country clusters.

As for potential stigma effects, these are found only in the case of Belgium. As can be seen, the interaction between lagged individual unemployment status and the cyclical unemployment rate has an associated coefficient of -0.192, statistically significant at 95%. This means that a decline (rise) in the unemployment rate is associated with an increase (decrease) in the impact that past unemployment status has on an individual, signalling potentially enhanced discrimination against individuals who became unemployed when the economy was growing. By contrast, in the cases of Bulgaria, Italy and Poland we obtain a positive coefficient for the same interaction, indicating that unemployment persistence behaves pro-cyclically – state dependence in unemployment actually increases (decreases) when the macroeconomic conditions

TABLE 2

Results of the RE probit models (equation 7) on youth unemployment state dependence and stigma effects, selected countries, 2004–2015

	y_{it-1}		$ue_{rp}y_{it-1}$	ue_{rp}
	α_1	*APE*	α_2	α_3
Austria	0.6707***	0.0912***	−0.1191	0.2167*
	[0.1385]	[0.0364]	[0.1281]	[0.1261]
Belgium	1.0638***	0.1803***	−0.1921**	0.0702
	[0.1271]	[0.0487]	[0.0814]	[0.0633]
Bulgaria	1.2782***	0.3291***	0.0814***	−0.017
	[0.0831]	[0.0406]	[0.0243]	[0.0938]
Czech Rep.	0.7738***	0.0503***	−0.0237	0.0146
	[0.0982]	[0.0179]	[0.0433]	[0.0499]
France	0.8951***	0.1400***	0.0013	−0.0088
	[0.0578]	[0.0209]	[0.0435]	[0.0274]
Greece	1.0886***	0.2340***	0.0083	0.0427*
	[0.0767]	[0.0320]	[0.0115]	[0.0256]
Hungary	0.5905***	0.0785***	0.0420*	0.1647***
	[0.0707]	[0.0171]	[0.0230]	[0.0500]
Italy	0.9662***	0.1968***	0.0896***	−0.0196
	[0.0432]	[0.0178]	[0.0147]	[0.0154]
Poland	1.0528***	0.2099***	0.0290**	0.0053
	[0.0485]	[0.0228]	[0.0133]	[0.0307]
Spain	0.6205***	0.1409***	0.0039	0.0132*
	[0.0419]	[0.0151]	[0.0076]	[0.0079]

Note: Results from estimating equation (7) using the longitudinal component of the EU-SILC from 2005 to 2015. Sample: Youth 17–29 years old. All specifications control for gender, age, age squared, maximum level of education attained, living outside the parental home, having a partner and the number of own children in the household. They also include year and region fixed effects.
***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

worsen (improve), result that cannot be attributed to stigmatization. For the rest of the countries analysed, no effect is found.

To confirm our findings, we ran a similar exercise as above, but instead of using the regional unemployment rate for the adult population, we took the youth unemployment rate at the regional level, as provided by Eurostat. The results can be found in Table A.3 of the Appendix. Genuine state dependence in youth unemployment is found to be positive and highly significant, and to occur at a similar level whether we use the total unemployment rate or the youth unemployment rate. As for stigma effects, the results are confirmed for Belgium (although significant at 90% confidence level instead of 95%), while Austria can be added as a country with a certain degree of stigmatization of young unemployed people. Thus, the estimates indicate the robustness of our results regarding the scarring effect of unemployment on young people, and provide further evidence of the negative association of this disadvantage over the business cycle when we consider tighter labour market conditions.

Following Biewen and Steffes (2010) and Omori (1997), we also computed our results by considering the unemployment rate at $t-1$ instead of at t, that is, following equation (10) above. The results are presented in Table A.4 of the Appendix. When past unemployment status is interacted with past unemployment risk, genuine state dependence is estimated at a similar level as in the previous analysis, but we do not find the presence of stigma effects in any of the countries analysed.

Are there differences by gender?

We present the results for gender-specific estimations in Table 3.[17] Columns (1)–(4) and (5)–(8) show the estimates for males and females respectively. Regarding the degree of genuine state dependence, we observe that past unemployment experiences ($t-1$) significantly increase the probability of being currently unemployed (t) for both genders.

Considering the magnitude of state dependence, the most affected are male youths living in Bulgaria, Greece and Poland, where the probability of being unemployed if previously unemployed increases by 32, 23 and 21 percentage points respectively. Next come Spain and France, with probabilities of around 14 percentage points. At the lower end of the distribution of the effect, we find the Czech Republic, with an estimate below 5 percentage points. In Belgium, we find much higher APEs associated with the state dependence effect among women than among men. Bulgaria stands out for having the most severe consequences for females, with an increased unemployment probability of 32 percentage points; it is followed by a group of countries including Belgium, Greece, Poland and Italy, which yield estimates of above 19 percentage points.

As for potential stigma effects, our results show evidence of these among both males and females in Belgium, with significant coefficients of −0.221 and −0.173 (albeit at 90% confidence level). There is further evidence that state dependence in unemployment decreases with unfavourable economic conditions in Austria, but only among females: a significant estimate of −0.321 (yet again at 90%). On the other hand, we find stronger effects of state dependence even when the economy is declining, especially in Italy, where it affects both males and females. Males in Bulgaria and females in Hungary show the same pattern.

The gender results presented above provide evidence for the existence of state dependence that affects both males and females. These effects are highly heterogeneous across the countries under analysis, and are stronger for females in some particular contexts. Concerning potential stigma effects, the evidence is weak.

How do the situations of young and of prime-aged people compare?

Individual characteristics like age are likely to affect differently labour market conditions such as job search intensity, mobility between jobs when looking for a

[17]The use of gender-specific regressions is justified based on the results of the corresponding Chow tests.

TABLE 3

Results of the RE probit models (equation 7) on youth unemployment state dependence and stigma effects by gender, selected countries, 2004–2015

	Male				Female			
	y_{it-1}		$ue_{rp}y_{it-1}$	ue_{rp}	y_{it-1}		$ue_{rp}y_{it-1}$	ue_{rp}
	α_1	*APE*	α_2	α_3	α_1	*APE*	α_2	α_3
Austria	0.6402***	0.0846*	0.1003	0.3468**	0.6846***	0.0866	−0.3210*	0.0481
	[0.1899]	[0.0466]	[0.1802]	[0.1752]	[0.2130]	[0.0537]	[0.1918]	[0.1882]
Belgium	0.8688***	0.1206**	−0.2212*	0.0447	1.2938***	0.2914***	−0.1735*	0.095
	[0.1850]	[0.0536]	[0.1260]	[0.1003]	[0.1711]	[0.0859]	[0.1036]	[0.0790]
Bulgaria	1.2581***	0.3176***	0.1145***	0.0297	1.3170***	0.3243***	0.0371	−0.0541
	[0.1111]	[0.0522]	[0.0329]	[0.1279]	[0.1241]	[0.0634]	[0.0367]	[0.1398]
Czech Rep.	0.7525***	0.0440**	−0.0704	0.039	0.8142***	0.0566*	0.0046	−0.008
	[0.1290]	[0.0212]	[0.0601]	[0.0711]	[0.1487]	[0.0293]	[0.0622]	[0.0697]
France	0.8704***	0.1342***	0.0824	−0.0239	0.9209***	0.1422***	−0.0818	0.0082
	[0.0815]	[0.0281]	[0.0608]	[0.0388]	[0.0823]	[0.0307]	[0.0628]	[0.0388]
Greece	1.0249***	0.2348***	0.0038	0.0639*	1.1696***	0.2592***	0.0161	0.0172
	[0.1030]	[0.0433]	[0.0159]	[0.0354]	[0.1153]	[0.0491]	[0.0165]	[0.0370]
Hungary	0.5495***	0.0718***	0.0163	0.2156***	0.6286***	0.0715***	0.0856**	0.1097
	[0.0925]	[0.0213]	[0.0297]	[0.0675]	[0.1107]	[0.0245]	[0.0369]	[0.0747]
Italy	0.9778***	0.1865***	0.0818***	−0.0049	0.9391***	0.1878***	0.0778***	−0.0316
	[0.0620]	[0.0240]	[0.0204]	[0.0226]	[0.0601]	[0.0251]	[0.0216]	[0.0210]
Poland	1.0416***	0.2049***	0.0287	0.0017	1.0094***	0.2000***	0.0294	0.0113
	[0.0706]	[0.0319]	[0.0197]	[0.0448]	[0.0698]	[0.0314]	[0.0185]	[0.0434]
Spain	0.6651***	0.1470***	−0.0033	0.016	0.5268***	0.1191***	0.011	0.0087
	[0.0593]	[0.0215]	[0.0107]	[0.0114]	[0.0619]	[0.0207]	[0.0111]	[0.0112]

Note: Results from estimating equation (7) using the longitudinal component of the EU-SILC from 2005 to 2015. Sample: Youth 17–29 years old. All specifications control for age, age squared, maximum level of education attained, living outside the parental home, having a partner and the number of own children in the household. They also include year and region fixed effects.
***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

suitable match, and gains (losses) in marginal productivity associated with human capital accumulation (skill depreciation). The association of all these characteristics with unemployment experiences at different phases of the business cycle might well affect the decisions of potential employers in the hiring process. Therefore, we now turn to an analysis of the state dependence and potential stigma effects among prime-aged individuals, between 30 and 44 years of age. Coefficient estimates are reported in Table A.5 of the Appendix.

The estimates on lagged unemployment shown in column (1) are all positive and statistically significant, indicating a strong positive association between unemployment in the previous period and the unemployment risk in the current period for mature workers. Comparing these results with those for younger individuals, presented in Table 2, we observe that the magnitude of the state dependence effect is greater for prime-aged workers in a handful of countries. This pattern is reflected in the APEs, showing that the effect is stronger for mature individuals, except in Austria, Belgium and Italy. Differences are particularly high in countries like France and the Czech Republic, where the probability of prime-aged workers being unemployed is above 20 and 10 percentage points respectively. Mature workers in France are more than 80% more likely than their younger counterparts to be unemployed, and are more than 100% more likely in the Czech Republic. Furthermore, Bulgaria and Greece show the strongest effects, with increased unemployment probabilities of around 34 and 29 percentage points, followed by France and Poland which are close to 24 percentage points.

Estimated results for the interaction between lagged individual unemployment and the cyclical unemployment rate are shown in column (3) of Table A.5. There is no significant evidence of stigma effects among prime-aged individuals in any of the countries under analysis. On the other hand, there is a group of countries, including Italy, Bulgaria, Poland, the Czech Republic, Greece and France that show positive and significant coefficients. This indicates that the disadvantage of having been unemployed in the previous period is smaller (larger) when unemployment is low (high).

Thus, our results show that the effect of genuine state dependence is considerably larger for prime-aged individuals in most countries. In order to check the robustness of this result, we have further replicated the analysis using the full age sample, including an indicator for prime-aged individuals, and the interaction of this indicator with our state dependence measure.[18] The results from these auxiliary regressions clearly indicate that prime-aged individuals are associated with lower levels of unemployment than are younger workers. Moreover, the coefficient of the interaction ($\alpha_1 y_{it-1} \times \text{prime} - \text{aged}$) is always positive and significant for all countries. Therefore, although the risk of being unemployed is lower among prime-aged individuals, the scarring effect of unemployment experiences in the previous period is substantially greater for them. In this scenario, younger workers might be associated with higher mobility and flexibility, and therefore may be less penalized by past unemployment

[18] Results are available from the authors upon request.

experiences. Young people are entering and leaving the labour market more often and more quickly than prime-aged individuals, and therefore, at an early stage of their labour market career, are less likely to accumulate long periods of unemployment – which is what our findings on unemployment state dependence reflect.[19]

Did the situation change for youth during the Great Recession?

In this section, we explore the extent to which unemployment experiences affected young workers before and after the Great Recession. Comparing the results that consider the precrisis period with those that refer to the years during and after the Great Recession will allow us to observe whether there are different patterns associated with state dependence and potential stigmatization that affected young individuals during these particular periods. Note that we can also now add results for the United Kingdom and Sweden; information at the regional level started to be provided by those two countries only in the 2007 (United Kingdom) and 2008 (Sweden) waves.

Columns (1)–(4) of Table 4 report the corresponding estimates for young individuals before the crisis (from 2004 to 2007) and columns (5)–(8) show the results associated with the Great Recession and its aftermath (from 2008 to 2015).[20] As in our main results, the coefficient on lagged unemployment for the period 2008–2015 is positive and statistically significant for all countries. Moreover, the estimates associated with the Great Recession are larger than those of the precrisis period in most countries; this provides evidence of the existence of stronger genuine state dependence in unemployment among youth during recession years. Importantly, the APEs associated with state dependence increase substantially after 2008 in both magnitude and significance. Note, in particular, the cases of Austria (from 4 to 13 percentage points), Belgium (from 8 to 27 percentage points) and Greece (from 12 to 35 percentage points).

As for stigma effects, we find robust evidence of these for the period 2004–2007 in Belgium and Poland, with statistically significant coefficients at 95% confidence level (−0.4621 and −0.0636 respectively). A negative coefficient is also found in Austria (−0.7933), but evidence of the stigmatization of young workers is rather weak, as it is only significant at 90%. During and after the Great Recession, the interaction between the cyclical unemployment risk and the individual past unemployment status (α_2) is found to be negative in Belgium, the Czech Republic and France – albeit at 90% only in the case of Belgium indicating that any stigma that past unemployment might have carried became weaker in this country during the period when unemployment affected a larger number of individuals. α_2 is found to be positive in Bulgaria (at 90%) and in Poland (at 95%), pointing to the increased difficulty of finding a job for those

[19]According to data from Eurostat, in 2018 in the European Union-28, the percentage of young people in long-term unemployment (12 months or more) was 21% (as a percentage of total unemployment). The figures for those in age groups 30–34, 35–39, 40–44 and 40–59 were 32.1%, 35.2%, 37.6% and 42.8% respectively.

[20]Note that, given the time span of our sample, we are only able to observe three transitions in this period; thus results must be interpreted with caution, particularly for the Czech Republic, Hungary and Poland, which joined the survey only in 2005 (see Table A.1 in the Appendix for specific details). We do not present results for Bulgaria, because it was not present in the sample until 2006.

TABLE 4

Results of the RE probit models (equation 7) on youth unemployment state dependence and stigma effects by period of time, selected countries, 2004–2015

	2004–2007				2008–2015			
	y_{it-1}		$ue_{rp}y_{it-1}$	ue_{rp}	y_{it-1}		$ue_{rp}y_{it-1}$	ue_{rp}
	α_1	*APE*	α_2	α_3	α_1	*APE*	α_1	*APE*
Austria	0.5112*	0.0374	−0.7933*	0.2034	0.7951***	0.1322***	−0.1112	0.2829**
	[0.3090]	[0.0490]	[0.4293]	[0.7456]	[0.1460]	[0.0481]	[0.1376]	[0.1357]
Belgium	0.6955**	0.0778	−0.4621**	0.0728	1.3100***	0.2710***	−0.1719*	0.0064
	[0.3250]	[0.0702]	[0.2013]	[0.2051]	[0.1229]	[0.0642]	[0.0936]	[0.0691]
Bulgaria	—	—	—	—	1.3002***	0.3123***	0.0487*	0.0757
					[0.0797]	[0.0347]	[0.0261]	[0.1218]
Czech Rep.	0.8563***	0.0504	−0.1264	0.0449	1.0260***	0.0932***	−0.1048**	0.0169
	[0.2996]	[0.0544]	[0.1000]	[0.1194]	[0.0934]	[0.0296]	[0.0501]	[0.0515]
France	0.7789***	0.0975**	0.2603**	−0.074	1.0317***	0.1999***	−0.1264**	0.0341
	[0.1698]	[0.0423]	[0.1033]	[0.0702]	[0.0620]	[0.0284]	[0.0505]	[0.0312]
Greece	0.8966***	0.1238**	−0.0137	0.1072	1.3056***	0.3511***	0.0082	0.0374
	[0.1906]	[0.0575]	[0.0399]	[0.1271]	[0.0724]	[0.0362]	[0.0114]	[0.0246]
Hungary	−0.233	0.0143***	0.3871	0.2813	0.7333***	0.1147***	0.0298	0.1987***
	[0.5541]	[0.0302]	[0.4090]	[0.2133]	[0.0711]	[0.0215]	[0.0233]	[0.0542]
Italy	1.1343***	0.2224***	0.2190***	−0.1190***	1.0171***	0.2371***	0.0099	0.0077
	[0.1025]	[0.0434]	[0.0328]	[0.0350]	[0.0443]	[0.0203]	[0.0170]	[0.0168]
Poland	0.7439***	0.1231**	−0.0636**	0.0396	1.2987***	0.2957***	0.0396**	−0.0106
	[0.1334]	[0.0481]	[0.0275]	[0.0634]	[0.0499]	[0.0271]	[0.0159]	[0.0355]
Spain	0.1995	0.0245	0.0293	0.0216	0.7180***	0.1818***	−0.0136	0.014
	[0.1295]	[0.0226]	[0.0269]	[0.0426]	[0.0430]	[0.0175]	[0.0084]	[0.0086]
Sweden					0.4071***	0.0637**	−0.067	0.178
					[0.1242]	[0.0311]	[0.1217]	[0.1673]
United Kingdom					0.5733***	0.0919**	−0.0729	0.0084
					[0.1715]	[0.0459]	[0.0812]	[0.1165]

Note: Results from estimating equation (7) using the longitudinal component of the EU-SILC from 2005 to 2015. Sample: Youth 17–29 years old. All specifications control for age, age squared, maximum level of education attained, living outside the parental home, having a partner and the number of own children in the household. They also include year and region fixed effects.
***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

unemployed at $t - 1$ when unemployment affected a larger share of the labour force in those two countries.

Furthermore, we need to consider that the Great Recession led to a substantial rise both in the unemployment rate and in the persistence of unemployment: these reached figures that were higher than their precrisis levels in most of the countries (OECD, 2016). Thus, there is concern regarding the possibility that high unemployment persistence might increase structural unemployment across the OECD countries (OECD, 2014, 2015). In a study of 28 OECD countries, Marques *et al*. (2017) find significant evidence of a structural break and hysteresis in the unemployment rates, indicating a change in the pattern of unemployment persistence after the Great Recession. Following on from that, we next consider a potential shift in the unemployment trend, by re-estimating our measure of the cyclical unemployment risk for two periods: 2004–2007 and 2008–2015. Thus, we test the robustness of our results, while reconsidering our previous assumption of a constant linear trend for the whole period of analysis. We report the main estimates of this exercise in Table A.6 in the Appendix.

Our findings regarding state dependence are robust to the use of a specific cyclical unemployment risk for each period, showing values in the coefficients and the corresponding APEs slightly smaller in magnitude for most of the countries. On the other hand, the results associated with the period 2004–2007 show evidence of potential stigma effects for Austria and Belgium, but not for Poland. Moreover, in the period 2008–2015, the interaction term between individual past unemployment status and the cyclical unemployment risk is no longer statistically significant in those countries where it was previously found. On the other hand, we find new evidence of stigma effects in Bulgaria, Poland and Spain. All in all, the results point to the fact that, while our results are robust when it comes to the scarring effect of unemployment, the same cannot be said for evidence of the processes of stigmatization of young workers in the European countries analysed.

Robustness checks

As was done in Biewen and Steffes (2010), we checked whether our results change when a wider definition of non-employment is used than one that refers strictly to unemployment. That way, not just the unemployed are considered, but also part-timers, marginally employed people and non-participants. In their case, Biewen and Steffes (2010) found that the evidence of stigma effects was stronger for Germany, suggesting that stigmatization was likely to occur partly through channels other than unemployment.

Following the same objective, we took data from the European Union – Labour Force Survey (EU-LFS) and computed at regional level three *proxies* of non-employment or under-employment: (definition 1) unemployed and inactive individuals, including those fulfilling domestic tasks and other inactive persons (except for students, the disabled and people in compulsory military service); (definition 2) unemployed individuals, together with individuals working in a part-time job because they could not find full-time employment, and individuals who wish to work more than the current number of hours; and finally (definition 3), adding to the definition of (2) also individuals who are seeking a new job (among other reasons, because they want to work more hours or fear that they will lose their current position).[21]

We matched these indicators with the EU-SILC data set used in the rest of the paper and found that state dependence in non-employment (or under-employment) is positive and statistically significant at 95% confidence level in all contexts and with coefficients at a level similar to the coefficients presented in Table 2 – see the detailed results in Tables A.7–A.9 in the Appendix. As for potential stigmatization, we find evidence of that only in Austria and the Czech Republic when we use an indicator of non-employment that considers the unemployed and some inactive individuals (definition 2). This means that in these particular contexts, the stigmatization of potential young workers also occurs through inactivity. No other relevant effect is found.

In addition, we considered whether our results differ when we account for previous labour market experience (Nordström Skans, 2011). We do so by taking advantage of a variable in the EU-SILC that collects data on the age at which an individual started his/her first job. We used this information to construct a binary indicator that takes the value 1 if an individual has some previous labour market experience, and 0 otherwise. We proceeded by interacting the lagged version of this indicator variable with the lagged labour market status indicator to determine whether the chances of young people exiting unemployment differ according to whether or not they have prior labour market experience. Moreover, we included an interaction between the stigma effect ($\mathrm{ue}_{rp} y_{i(t-1)}$) and this indicator variable. The results (available from the authors on request) indicate that previous labour market experience reduces the scarring effect of unemployment in only 5 of the 10 countries analysed, while no distinctive effect is found in the case of stigmatization.[22]

Finally, we tested the sensitivity of our findings on stigmatization by considering variations in the cyclical unemployment rate as they apply to all countries. Thus, we calculated the APEs associated with α_2 for three values of the cyclical unemployment rate risk (−2, 0, +2). The results can be found in Table A.10 in the Appendix. In those countries where α_2 is not statistically significant, different values of the unemployment rate risk are associated with a similar probability of being unemployed because of previous unemployment. The same is not true of countries with a positive and significant effect for α_2: the risk of being unemployed because of previous unemployment increases with growing unemployment. In the case of Belgium – the only country where we could establish stigmatization of workers who have previously been unemployed – we can confirm that lower values of the unemployment rate risk increase the probability of individuals being unemployed because of previous unemployment; this confirms the fact that employers are reluctant to hire someone who became unemployed while the economy was growing.

[21]Note that unlike in Biewen and Steffes (2010) we are unable to change our dependent variable following these definitions, because in the EU-SILC data set there is no information, for example, on whether part-time workers wish to work more hours or whether those employed are looking for another job because they fear they will lose their current position. Thus, this exercise simply measures the robustness of our results to *proxies* of the business cycle that go beyond the unemployment rate.

[22]We would like to thank a referee for suggesting this analysis.

Unemployment state dependence and institutional settings

Our study finds considerable heterogeneity in the scarring effect of unemployment among young workers in different European countries. As these countries also differ substantially with respect to institutional and cultural settings, we investigate whether there is any association between unemployment state dependence and some of these factors.[23] Our analysis is based on just a few observations, and hence any possible explanation has, at most, a tentative character. In particular, we consider: (i) the OECD Employment Protection Index (differentiated for permanent and temporary contracts), (ii) the unemployment rate, (iii) union density, (iv) benefit duration and (v) the tax wedge rate (for a detailed description of each indicator, see Appendix B and Table A.11 for descriptive statistics by country). We use the mean value of each indicator for the period 2008–2013 provided by the OECD database (which excludes Bulgaria) and the APEs shown in Table 4 for the period 2008–2015, as this enables us to include the largest number of countries. The respective scatter plots are presented in Figure A.2 in the Appendix, and the fitted line shows the direction of the association.

We find a positive correlation between the strictness of employment protection (both for permanent and for temporary workers) and the scarring effect of unemployment. Such a result indicates that employers are more reluctant to hire when the burden for dismissal is elevated. For the unemployment rate, we also find a positive relationship. However, after removing the two outliers, Greece and Spain, the relationship becomes less clear. A higher level of union density seems to lower persistence in unemployment marginally, but the association is weak. A negative association is found with respect to benefit duration. This finding is counter-intuitive as, in general, a more generous benefit system is often associated with weaker incentives to leave the ranks of the unemployed. However, two countries (Greece and Italy) seem to have a substantial impact on this association. Finally, we find a (weak) positive association with the tax wedge rate; this is in line with our expectations, as higher tax rates might disincentivize entry to employment.

VI. Conclusions

This paper studies unemployment persistence among young people in Europe in the context of the Great Recession. To that end, we have estimated the extent to which the probability of being unemployed during the previous year in itself influences the probability of being currently unemployed; that is, we provide a measure of *genuine state dependence* that accounts for observed and unobserved heterogeneity and the initial conditions problem. Moreover, we consider whether the young unemployed are being stigmatized in today's labour market, in the sense of being even more discriminated against if they became unemployed when the economy was growing. We use data from all the available waves of the EU-SILC, which runs from 2004 to 2015, and for a selection of 12 countries.

[23]In general, the economic literature (e.g. Blanchard and Wolfers, 2000; Belot and van Ours, 2004) has found that institutional and cultural factors have an impact on labour market outcomes.

Our main findings indicate that young people suffer from an important degree of genuine state dependence in unemployment, but that the effect varies greatly across Europe. Of the countries analysed, it is in Bulgaria, Greece and Poland that the effect is strongest. For instance, in Bulgaria, the probability of being unemployed at t increases by 32 percentage points if the person was unemployed at $t-1$, compared to someone employed or inactive at $t-1$. The effect is lowest in the Czech Republic. The variety in the magnitude of the effects found prevents the establishment of clusters of countries, either by geographical region or by type of welfare state regime.

When looking at the results by gender, the effects are somehow stronger for females than for males. Comparison of our results for young people with those for prime-aged individuals indicates that in fact, in the large majority of countries, it is rather the mature unemployed that suffer a higher degree of genuine state dependence. Additional results indicate that while prime-aged individuals are associated with overall lower levels of unemployment, when they are unemployed they suffer a higher degree of unemployment persistence than young people. The fact that young individuals are less affected by the scarring effect of unemployment is consistent with previous results in the literature and is associated with ‘job shopping’ behaviour or the propensity to change jobs often during youth (Arulampalam *et al*., 2000). Moreover, we have compared the results for the period 2004–2007 with those for 2008–2015 to observe that during and after the Great Recession the genuine state dependence effect increased in magnitude for young people. That is, the likelihood of being unemployed because of previous unemployment worsened for those between 17 and 29 years of age in the context of the economic crisis.

In relation to potential stigma effects suffered by the young unemployed during the period, the empirical evidence found for Europe is rather weak. As a matter of fact, we have only found such effects in the case of Belgium when considering the whole period of analysis. This effect is shared similarly by males and females in that country. At this point, it is worth noting that the particularly generous welfare state regime in Belgium, with its high minimum wages and unemployment benefits of unlimited duration, has been shown by the literature to induce more unemployment and wage penalties, reinforcing persistence especially among low-educated workers and particularly during recessions (Cockx and Ghirelli, 2016). In the rest of the countries, either no effect is found or the effect is positive, indicating that state dependence in unemployment actually increases when the unemployment rate is rising, thus contributing to the increased difficulties young unemployed people have in finding a job – but also pointing to the fact that the youth are not necessarily discriminated against because they became unemployed when the economy was growing. Importantly, when we restrict our analysis to the period from 2008 to 2015, when the unemployment rate was mostly above its trend, we learn that only in the Czech Republic and in France did any stigma that past unemployment might carry become stronger, whereas the effect vanishes in Austria and Poland and decreases in Belgium. Hence, we interpret these results as evidence of weaker stigma effects during the Great Recession.

This paper suffers from a number of limitations. First, our analysis could be carried out only for an arbitrary selection of 12 European countries: those that in the EU-SILC provided information at the regional level. This implies that our results are not representative of the experiences of young people across Europe. Second, due to the design of the longitudinal component of the EU-SILC, only three transitions between labour market statuses per individual could be observed. Therefore, our findings are relevant as short-term consequences of the scarring effect of unemployment. Finally, our analysis is based on a period of time that generally saw an increase in the unemployment rate; thus, future analysis should account for a longer period of time that would contain more phases of the business cycle.

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Supporting Information

Additional Supporting Information may be found in the online version of this article:

Appendix A. Additional tables and figures.
Appendix B. Unemployment state dependence and institutional settings (definitions).