EUROPEAN CENTRAL BANK

EUROSYSTEM

Working Paper Series

Corinne Petrakis, David Sondermann

Hidden in the aggregate: the cyclicality of EU labour force participation

No 3275

Abstract

We estimate the cyclical response of labour force participation to growth shocks across EU regions. We use a shift-share instrumental variable approach and local projections on EU Labour Force Survey microdata covering 15 countries and 114 regions over 2000–2020. The aggregate labour force participation rate is remarkably resilient, a clear contrast to the highly cyclical participation rate documented for the U.S. However, this resilience masks pronounced demographic heterogeneity. Men's participation declines significantly while women's remains stable, reflecting sectoral segregation and the added worker effect. Young workers exhibit the most persistent responses, with effects lasting up to eight years, consistent with hysteresis through human capital erosion. Less-educated and non-native workers bear disproportionate employment costs. The findings imply that aggregate slack measures may miss important distributional dimensions relevant for monetary, fiscal, and structural policy.

JEL Classification: J21, E24, E32, R23

Keywords: labour force participation rate, Bartik instrument, demographic heterogeneity, local projections.

Non-technical summary

How does the labour force respond to economic downturns? When GDP declines, some people lose their jobs and become unemployed, while others may stop looking for work altogether and leave the labour force entirely. Whether people stay in or drop out of the labour force during recessions matters greatly for policymakers: if participation declines alongside employment, the unemployment rate understates the true extent of slack in the economy, with consequences for monetary policy, fiscal revenues, and the design of social safety nets.

For the United States, recent research has shown that the labour force participation rate is highly cyclical. It declines for about four years after a negative shock and takes roughly eight years to fully recover.

Whether the same holds for the EU is an open question, given Europe's stronger employment protection, more generous welfare systems, and greater institutional diversity.

This paper provides the first causal estimate of this question for the EU. Using microdata from the EU Labour Force Survey covering 15 countries and 114 regions over 2000–2020, we trace how the labour force participation rate, the employment rate, and the unemployment rate respond to regional growth shocks. To isolate genuinely exogenous shocks, we construct an instrumental variable that exploits differences in industrial structure across EU regions.

Our main finding is that the aggregate EU labour force participation rate is remarkably resilient to growth shocks. While employment falls significantly and unemployment rises persistently, with effects lasting four to five years, participation barely moves. This stands in sharp contrast to the U.S. experience.

However, this aggregate stability masks substantial differences across groups. Men's participation declines significantly after a negative shock, while women's remains stable, reflecting both the concentration of men in cyclically sensitive sectors like construction and manufacturing, and the tendency of women to enter the labour force when their partner loses income (the so-called "added worker effect"). Young workers (aged 15–24) are hit hardest: their participation falls and does not recover for up to eight years, suggesting that recessions can long-lastingly damage an entire generation's attachment to the labour market. Less-educated and non-native workers also bear disproportionate costs.

These findings carry several policy implications. For monetary policy, the resilience of aggregate participation suggests that the unemployment rate captures cyclical slack in the EU more reliably than in the U.S., but the increase in people who want to work yet have stopped searching points to hidden slack beneath the surface. For fiscal and structural policy, the long-lasting effects on youth participation make a strong case for early intervention through training programmes, employment subsidies, and apprenticeship schemes that prevent temporary job losses from becoming permanent exits. For social policy, the gendered pattern of adjustment, where women absorb household income shocks by entering the labour force, underscores the importance of childcare provision and flexible work arrangements.

More broadly, the finding that aggregate resilience is the product of offsetting vulnerabilities across groups, rather than uniform stability, suggests that targeted policies are essential to complement the economy's natural shock-absorbing capacity.

1 Introduction

Understanding how labour markets adjust to business cycle fluctuations is central to the conduct of monetary policy and to the design of structural policies. The cyclical behaviour of the labour force participation rate (LFPR) determines the extent to which slack in the economy is captured by conventional indicators such as the unemployment rate. If participation responds to growth shocks, the unemployment rate alone provides an incomplete picture of labour market tightness and, by extension, of the inflationary pressures relevant for monetary policy (Blanchard and Summers, 1986). Beyond the central bank, LFPR cyclicality carries direct implications for fiscal policies (through its effect on the tax base and social expenditure), for active labour market policies (through the identification of groups at risk of permanent withdrawal), and for social policy more broadly (through its distributional consequences across demographic groups).

In the United States (U.S.), Cajner et al. (2025) document that the LFPR moves with the cycle, with responses that substantially outlast those of the unemployment rate. After a negative output shock, the U.S. LFPR declines for roughly four years and takes about eight years to fully recover. This finding has significant implications: observers who focus solely on the unemployment rate may underestimate the extent of remaining slack in the labour market for several years after a recession. Does the same hold true for the EU? Several features of European labour markets, including stronger employment protection, more generous welfare systems, short-time work schemes, and substantial cross-country institutional heterogeneity, could produce fundamentally different participation dynamics. The LFPR in the EU moved significantly in recent years, explaining a significant part of employment dynamics (Arce and Sondermann, 2026). But is this a cyclical development or is the LFPR mainly driven by persistent structural trends, such as the increase in female and older participation (Berson and Botelho, 2023)? A systematic analysis of European LFPR dynamics, separating between structural and cyclical movements, is so far missing. We aim to fill this gap.

Our paper contributes to the literature on European labour market dynamics in several ways. First, it is the first study to provide causal estimates of LFPR cyclicality across EU regions using exogenous output shocks. Cajner et al. (2025) do so for the U.S., but their findings cannot be extrapolated to the EU given the fundamental differences in labour market institutions. Our paper is, to our knowledge, the first to apply the full Cajner et al. (2025) framework, output-based Bartik instruments, local projections, and age-sex adjustment, to the

EU, enabling a direct transatlantic comparison using a comparable methodology. Second, by examining all major demographic dimensions (gender, age, education, citizenship, industry) within a single econometric framework, we can trace the sources of aggregate resilience to specific offsetting subgroup dynamics, something that the existing literature, which typically focuses on one dimension at a time, cannot do. Third, our leave-one-country-out construction of the Bartik instrument is a methodological adaptation to the EU's institutional structure that addresses the high within-country spatial integration absent in the U.S. Fourth, we connect the demographic heterogeneity in LFPR responses to the hysteresis literature, providing new evidence on the long-run consequences of demand shocks for youth participation. Finally, our findings speak to the recent policy debate on EU labour market resilience, where participation dynamics have come to the fore in the context of tightening labour markets and demographic change (Galí, 2020).

We employ the EU Labour Force Survey (LFS) microdata covering 15 countries and 114 NUTS-2 regions over the period 2000–2020. Following Cajner et al. (2025), we estimate impulse response functions (IRFs) of the LFPR, employment rate, and the unemployment rate to regional growth shocks, using local projections (à la Jordà, 2005). To address the endogeneity of GDP, we construct a shift-share (Bartik, 1991) instrument based on leave-one-country-out industry-level GVA growth rates. We adjust our outcome variables for changes in the age-sex composition of the population, thereby isolating cyclical responses from structural demographic shifts such as population ageing and rising female participation.

Our main finding is that the EU LFPR is remarkably resilient to growth shocks. While the employment rate and unemployment rate respond strongly and persistently, with effects lasting four to five years, the aggregate LFPR remains statistically insignificant at all horizons. This stands in sharp contrast to the U.S. results of Cajner et al. (2025), where the LFPR responds significantly and with long-lived effects. However, this aggregate resilience conceals substantial heterogeneity across demographic groups. Men's LFPR declines significantly following negative shocks, whereas women's participation remains stable, consistent with industry-level gender segregation and the added worker effect. Among age groups, the young (15–24) exhibit the largest and most persistent responses, with LFPR effects lasting up to eight years, pointing to potential scarring effects. Less-educated and non-native workers are also disproportionately affected.

These results carry direct implications across policy domains. For central banks, the finding that aggregate LFPR resilience in the EU coexists with highly heterogeneous subgroup responses implies that standard measures of labour market slack may adequately capture aggregate cyclical conditions but miss distributional dimensions that matter for wage formation and price dynamics across sectors. For fiscal and structural policy, the persistent LFPR decline among young workers following growth shocks underscores the importance of active labour market policies to avoid hysteresis. The pronounced vulnerability of non-native and less-educated workers further highlights the need for targeted integration programmes and social safety nets during downturns. For social policy, the gendered nature of the response, with the added worker effect operating through women entering the labour force when men lose jobs, has implications for the design of childcare support, flexible work arrangements, and gender equality policies.

The remainder of the paper is organised as follows. Section 2 reviews the related literature. Section 3 presents the econometric framework and identification strategy. Section 4 describes the data. Section 5 presents results, including the aggregate IRFs, demographic heterogeneity, and industry-level analysis. Section 6 discusses robustness checks and instrument validity. Section 7 concludes.

2 Related Literature

Our paper relates to three main strands of the literature: (1) the cyclicality of labour force participation, (2) the heterogeneous labour market effects of shocks across demographic groups, and (3) the use of shift-share instruments for causal identification.

LFPR cyclicality and labour market adjustment. The foundational work of Blanchard and Katz (1992) established that negative employment shocks at the U.S. state level generate persistent responses in employment but transitory effects on unemployment and participation, with migration serving as the primary adjustment mechanism. Cajner et al. (2025) revisit this question using output shocks rather than employment shocks, finding that the U.S. LFPR is in fact highly cyclical, with responses that substantially outlast those of the unemployment rate. The LFPR declines for about four years after a negative shock and takes approximately eight years to fully recover, with the decline driven largely by individuals with home and family responsibilities and younger individuals spending time in school. Our paper applies their empirical framework to the EU, where labour mobility is more limited and institutional structures differ markedly. Decressin and Fatás (1995) provide the classic EU–U.S. comparison of regional labour market adjustment but focus on employment rather than participation and use data predating the euro. Our paper updates and extends this comparison using modern causal methods focused specifically on the participation margin (see also Dao et al., 2017, for a recent US regional adjustment analysis). Bräuer and Kersting (2023) provide a complementary perspective, showing that the persistence of local employment shocks depends critically on the educational composition of the affected region, counties with higher shares of college-educated adults experience significantly less persistent unemployment.

Several mechanisms underpin LFPR cyclicality. The discouraged worker effect, whereby deteriorating job prospects lead individuals to exit the labour force, has been documented as the dominant force in OECD countries (Lee and Parasnis, 2014). Evans et al. (2019) use Australian data to confirm that cyclical movements in participation depend on economic conditions, with younger workers, prime-age females, and older males being most responsive. Counteracting the discouraged worker effect, the added worker effect, whereby secondary earners enter the labour force when the primary earner loses income, has been shown to vary significantly across European welfare state regimes. Bredtmann et al. (2017) find strong added worker effects in Mediterranean countries with low welfare generosity, while Hardoy and Schøne (2014) and Cammeraat et al. (2023) find negligible effects in Nordic and Continental European countries where state insurance substitutes for spousal labour supply. Tuzemen and Van Zandweghe (2018) provide a structural perspective, showing that technological shocks can generate countercyclical LFPR responses while discount factor shocks produce procyclical ones, with the latter dominating in the data.

Heterogeneous effects across demographic groups. A robust finding in the literature is that recessions affect labour market participants unevenly. Hoynes et al. (2012) document that less-educated workers, men, and minorities bear disproportionate costs during U.S. recessions. For the EU, Verdugo and Allègre (2020) show that gender segregation across industries played a key role during the Great Recession: regions with sharper declines in male-dominated sectors saw larger increases in female LFPR. They argue that discerning demand shocks by gender is essential for understanding participation disparities between men and women. More broadly, the secular rise in female LFPR has been attributed to changes in social norms (Fernández, 2013), European employment policies supporting gender equality (León, 2009), and declining birth rates that enabled women to invest more in careers (Cipollone et al., 2014; Vlasblom and Schippers, 2004).

Bell and Blanchflower (2011) highlight the particular vulnerability of young workers, who face higher separation rates, lower job search capital, and seniority-weighted redundancy systems that make them cheaper to dismiss. They document a negative correlation between youth unemployment and participation, suggesting a strong discouraged worker effect. Younger workers often retreat into education during downturns (Dellas and Koubi, 2003), which may partly explain their longer-lived LFPR responses. For non-native workers, the literature documents an "ethnic penalty", a lower probability of employment compared to natives even after controlling for observable characteristics, that is amplified during recessions (Reyneri and Fullin, 2011; Bratsberg et al., 2016). Bound and Holzer (2000) provide further evidence using U.S. metropolitan data from the 1980s, finding that the less educated, the young, and Black workers were most adversely affected by local demand shocks, with migration only partially alleviating these effects.

The unprecedented shocks of the past few years have provided further evidence on these dynamics. Béland et al. (2020) document sharp declines in hours worked and LFPR during the COVID-19 pandemic, with less-educated and male workers being more susceptible. Pouliakas and Branka (2020) confirm that workers in occupations requiring close physical proximity were disproportionately affected. The rapid recovery of European labour markets after the pandemic, driven in part by labour hoarding, short-time work schemes, and strong labour demand, has renewed interest in understanding the structural foundations of EU labour market resilience (Balleer et al., 2009; Bodnár, 2018).

Shift-share instruments. We follow a growing literature employing Bartik-type instruments to identify exogenous variation in local economic conditions (Bartik, 1991; Autor et al., 2013; Goldsmith-Pinkham et al., 2020; Notowidigdo, 2020). The identification assumptions underlying these instruments have been formalised by Goldsmith-Pinkham et al. (2020), who emphasise the exogeneity of the shares, and by Borusyak et al. (2022), who emphasise the quasi-random assignment of shocks. Our leave-one-country-out construction of GVA growth rates ensures that the industry-level shocks are uncorrelated with region-specific labour market conditions, aligning most closely with the shock-level exogeneity framework. The use of lagged industry shares further mitigates concerns about endogenous industrial composition.

3 Econometric Framework

We employ a local projection framework to trace the dynamic response of the LFPR, the employment rate, and the unemployment rate to business cycle shocks. Section 3.1 sets out the baseline specification. Three methodological choices are central to our identification strategy. First, we use output growth rather than employment changes as our shock variable, which captures both the employment and productivity channels of the business cycle (Section 3.2). Second, we instrument regional output growth with a Bartik-type shift-share variable to address endogeneity, adapting the standard approach to the EU's institutional structure through a leave-one-country-out construction (Section 3.3). Third, we adjust our outcome variables for changes in the age-sex composition of the population, isolating cyclical responses from structural trends such as ageing and rising female participation (Section 3.4).

3.1 Local Projections

We estimate the dynamic causal effect of growth shocks on labour market outcomes using local projections (Jordà, 2005). For outcome variable y (LFPR, employment or unemployment rate) in region r at time t, we estimate:

$$y_{r,t+k} - y_{r,t-1} = \beta^{(k)} Shock_{r,t} + \Theta W_{r,t} + \varepsilon_{r,t+k} \tag{1}$$

where k indexes the horizon ahead, $\text{Shock}_{r,t}$ is the business cycle shock measured as the percentage point change in GDP growth, and $W_{r,t}$ includes regional and year fixed effects as well as region-specific time trends.

While the age-sex adjustment (to be explained in Section 3.4) removes variation driven by shifts in the demographic composition of the regional population, it does not eliminate within-cell secular trends, for instance, rising female participation within a given age group driven by evolving social norms, expanding educational attainment, or policy reforms such as changes in pension eligibility or childcare provision. Our panel specification addresses this concern through three layers of controls, as included in Equation 1. Year fixed effects absorb any trend or shock common to all regions in a given year, including EU-wide secular increases in participation. Region fixed effects remove time-invariant level differences across regions. Crucially, the inclusion of region-specific time trends allows each region to follow its own secular trajectory, ensuring that identification comes from deviations of outcomes from each region's idiosyncratic trend path, net of common time effects. This structure prevents the confounding of cyclical responses with heterogeneous structural dynamics that would arise if, for example, regions experiencing larger growth shocks also happened to be undergoing faster structural increases in participation due to concurrent institutional reforms.

The inclusion of region-specific linear time trends is a strong control, and its implications for identification warrant explicit discussion. Identification comes from deviations of outcomes from each region's linear trajectory, net of common year effects. Three implications follow. First, secular changes that unfold approximately linearly over the sample (gradual female labour-market integration, phased-in pension reforms, the secular rise in educational attainment) are absorbed by the trends and do not contaminate our cyclical estimates. Second, the trends shape the interpretation of persistent coefficients: a response that remains negative at long horizons is identified as a shortfall relative to the region's own pre-shock secular path, not as a shift in that path itself. When we describe persistent responses as scarring (Section 5.3.2), it is in this precise sense (a medium-run deviation from trend that has not reverted within the horizon we estimate) rather than as a claim about permanent changes in long-run equilibrium participation, which a specification with linear trends cannot separately identify. Third, a natural concern is that region-specific trends might themselves absorb part of a slow-moving cyclical response, biasing persistence estimates towards zero. Section 6.4 addresses this directly: omitting the trends leads to a loss of significance rather than larger persistent effects, while adding quadratic trends leaves the results unchanged, indicating that the trends sharpen identification rather than absorb the responses of interest.

The coefficient $\beta^{(k)}$ traces the impulse response function at each horizon k. We also show pre-trends from horizon -3 onwards, normalising with respect to horizon -1, to verify the absence of pre-existing trends in our outcome variables.

We cluster standard errors at the NUTS-2 regional level. This is the level at which the outcome, regressor, and instrument are defined, and the level at which within-unit serial correlation across local-projection horizons is most evident. For local projections specifically, Montiel Olea and Plagborg-Møller (2021) show that heteroskedasticity-robust standard errors deliver asymptotically valid inference without additional corrections for serial dependence; Li et al. (2024) corroborate the robustness properties of LP-based inference across a wide range of data-generating processes. Clustering at the panel cross-section unit additionally accommodates the mechanical serial correlation induced by the overlapping nature of the dependent variable across horizons

k.

Three further sources of dependence in our setting warrant attention and motivate the robustness exercises in Section 6.4. First, NUTS-2 regions are nested within countries that share fiscal frameworks and labour-market institutions, and for euro-area members a common monetary policy stance. Second, residuals may be correlated across regions within the same calendar year for reasons not absorbed by year fixed effects. Third, by construction the shift-share instrument exposes all regions to common industry-level shocks, weighted by region-specific shares, the form of dependence highlighted by Adão et al. (2019). We assess sensitivity to each of these by reporting standard errors clustered at the country level, two-way clustered by region and year, and computed using the exposure-robust procedure of Adão et al. (2019) designed specifically for shift-share IV.

Each regression is weighted by the population share of the relevant demographic group, ensuring that our estimates are representative at the EU level and that regions with very few observations do not disproportionately influence the results.

We prefer local projections over vector autoregressions (VARs) for two reasons. First, LPs do not require the assumption of a correctly specified data-generating process, they are robust to misspecification of the lag structure and do not require invertibility. Second, LPs straightforwardly accommodate the instrumental variable approach described below.

3.2 Growth Shocks vs. Employment Shocks

Our business cycle shock is in principle defined as the percentage change in regional output growth from one year to the next. However, we will discuss in Section 3.3 the need to instrument it. Following Cajner et al. (2025), we prefer output shocks over employment-based measures for three reasons. First, changes in unemployment overlook productivity fluctuations: a negative output shock may lower both employment and output per worker, and an employment-based approach would miss the latter channel. Second, there may be a delay between a business cycle shock and its transmission to the labour market, so that output changes capture the timing of the shock more precisely. Third, estimating IRFs of our outcome variables to output shocks is more naturally interpretable in terms of Okun's law, a central concept in the literature on cyclical labour market dynamics.

In Section 6.1, we confirm that productivity accounts for a substantial share of the output shocks identified by our instrument, validating this methodological choice.

3.3 Bartik Instrument

Output growth at the regional level is likely endogenous: employment and output are jointly determined, and productivity shocks may affect both. We instrument regional GDP growth with a Bartik (shift-share) variable constructed as:

$$Z_{r,t} = \sum_q \Delta \mathrm{GVA}_{q,-c,t} \times \omega_{q,r,t-5} \tag{2}$$

where $\Delta\mathrm{GVA}_{q,-c,t}$ is the growth rate of gross value added in industry q at time t, computed across all regions *excluding* those in the country c where region r is located ("leave-one-country-out approach"), and $\omega_{q,r,t-5}$ is the three-year moving average of industry q's share of total GDP in region r, lagged five years.

The leave-one-country-out construction is essential in the European context. Unlike U.S. states, regions within the same EU country share fiscal policy, labour market institutions, regulatory frameworks, and a high degree of economic integration. As a result, regional economic conditions are likely to be considerably more correlated within countries than across countries. To demonstrate this, Table 1 reports the average pairwise correlation of annual NUTS-2 GDP growth for region pairs located within the same country and across different countries. We consider three specifications: (i) the full sample (2000–2020), (ii) the pre-pandemic sample (2000–2019), and (iii) detrended GDP growth, obtained after removing year fixed effects to isolate country-specific co-movement from common European shocks. Across all specifications, within-country correlations are substantially higher than across-country correlations. Most notably, after removing common year effects, the average correlation across countries is close to zero, whereas substantial within-country co-movement remains. These results indicate that regions within the same country continue to share important country-specific shocks even after accounting for common macroeconomic fluctuations. Consequently, constructing the Bartik instrument using national or EU-wide industry growth would risk contaminating the shift component with country-specific shocks that are correlated with regional labour market outcomes. Excluding regions from the same country ensures that the shift component captures only foreign industry-level developments, strengthening the credibility of the identification strategy.

The instrument must satisfy three conditions: (i) relevance, industry-level shocks abroad predict regional GDP growth through the pre-existing industry mix; (ii) contemporaneous exogeneity, the leave-out construction ensures that industry shocks are uncorrelated with region-specific changes in labour market outcomes; and (iii) lead-lag exogeneity, no correlation between the instrument and future or past outcome innovations.

Measure	Pair Type	Mean Correlation	Pair Count	Mean Overlap (years)
Detrended, year FE	Across country	-0.04	5349	18
	Within country	0.62	537	20
Raw, 2000-2019	Across country	0.45	5349	17
	Within country	0.77	537	19
Raw, 2000-2020	Across country	0.55	5349	18
	Within country	0.83	537	20

Table 1: Pairwise correlations for regions within and across country. We consider three different cases; 1. detrended data, controlling for year fixed effects 2. Raw data from 2000-2019 (where we exclude the Covid-19 pandemic) and 3. Raw data, 2000-2020.

The first condition is likely to be satisfied given that changes occurring to an industry on a national level due to e.g. changes in technology or demand, will not only impact national output, but regional output as well. Using the "leave-out-one" approach to construct our instrument, ensures that any industry shocks are uncorrelated to regional changes in labour market outcomes, and therefore condition (ii) is met. Our instrument also needs to satisfy lead-lag exogeneity. In Section 6 we demonstrate that indeed, our instrument satisfies this condition as well and hence constitutes an appropriate instrument, capturing exogenous growth shocks. Therefore, we follow a growing literature that employs Bartik instruments to identify exogenous shock for the purpose of studying their effects on labour market dynamics (see Autor et al. (2013), Goldsmith-Pinkham et al. (2020) and Notowidigdo (2020)).

A final clarification concerns what our design identifies. EU industry-level cycles co-move across countries, so part of the foreign shift in Equation (2) is common to all regions in a given year. This common component, together with any EU-wide disturbance that moves participation in all regions alike, such as an area-wide monetary or global demand impulse, is absorbed by the year fixed effects in Equation (1). Identification of $\beta^{(k)}$ therefore rests on the *differential* exposure of regions to foreign industry shocks through their pre-period industrial structure: regions whose historical industry mix loads more heavily on currently contracting foreign industries receive larger instrumented growth shocks, and it is the labour-market response along this exposure dimension that we estimate. This is a general property of shift-share designs rather than a limitation specific to our setting (Borusyak et al., 2022; Goldsmith-Pinkham et al., 2020). It also delimits the interpretation of our estimates: they describe how participation responds to growth shocks transmitted *differentially* across regions, the object relevant for the transatlantic comparison with Cajner et al. (2025), whose identification rests on the same differential variation, whereas movements common to all EU regions in a given year, whether cyclical or secular, are differenced out and are not estimated here.

3.4 Age-Sex Adjustment

Structural demographic shifts, population ageing or changing gender composition, can confound the estimation of cyclical responses. We follow Cajner et al. (2025) and adjust outcome variables for changes in the age-sex composition of the population. Specifically, for individual i in region r at time t, we regress the outcome variable on age-sex fixed effects:

$$y_{i,r,t} = \theta_{\text{age}(i),\text{sex}(i)} + \tilde{y}_{i,r,t} \tag{3}$$

Using the residuals $\tilde{y}_{i,r,t}$, aggregated to the regional level with population weights, we obtain the age-sex adjusted outcome variable for each region-year. This ensures that our IRFs capture within-group cyclical responses rather than compositional shifts. One concern is that this adjustment may not account for within-cell structural drift (driven by changes in education, policy reforms, or cultural shifts within age-sex cells). We address this by showing that additional adjustment for education yields virtually identical results, suggesting that the age-sex partition captures the dominant sources of compositional variation.

4 Data

We draw on three data sources from Eurostat: the EU Labour Force Survey (LFS) microdata, regional GDP at the NUTS-2 level, and gross value added (GVA) by industry and region. We construct the Bartik instrument using 19 countries [1] and 142 regions. Due to data availability of the outcome variables for some countries, we run our regressions, using 15 of those countries [2], 114 regions, and 10 non-overlapping industries over the period 2000–2020[3]. Including the

[1] Austria, Belgium, Croatia, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Lithuania, Poland, Portugal, Slovakia, Slovenia, Spain and Sweden.

[2] The 15 countries are Austria, Belgium, Croatia, Denmark, Finland, France, Germany, Greece, Hungary, Italy, Lithuania, Portugal, Slovenia, Spain and Sweden.

[3] 10 industries are included in the sample as in Eurostat; A: Agriculture, forestry and fishing; B,C,D,E: Mining, Energy, Manufacturing and Water supply; F: Construction; G-I: Domestic trade, transport, accommodation and

additional countries in the construction of the Bartik shift increases the external variation used to measure foreign industry-level shocks and strengthens the leave-one-country-out design by reducing exposure to country-specific shocks from the estimation sample [4].

The LFS provides individual-level information on labour force status, demographics, and other characteristics. We use these microdata to compute the LFPR, employment rate, and unemployment rate at the regional level, applying survey weights to ensure population representativeness. For the demographic subgroup analysis, we condition on gender, age, educational attainment, citizenship status, and industry of employment. The LFS follows a rotating panel design whereby each household is surveyed for five consecutive quarters before being replaced. Because our dataset follows an annual time series, we cannot track individual households over time, which precludes the analysis of inflows and outflows between labour market states. We return to this limitation in the conclusion.

A note on terminology is warranted, since migration enters our analysis in two conceptually distinct ways. First, *citizenship status* (native, EU citizen, non-EU citizen) is a demographic attribute of individuals in the LFS resident-population sample, which we use as a dimension of the heterogeneity analysis in Section 5.3.4. It captures migration background rather than migration flows. Second, *migration flows*, whether across regions within a country, across EU borders, or into and out of the EU, alter the composition of the regional resident population over time. The LFS samples the population resident in a region at the time of the survey: individuals who leave a region exit both the numerator (labour force) and the denominator (working-age population) of that region's LFPR, while new arrivals enter both. Regional participation rates therefore reflect the behaviour of the population resident in each period and are not invariant to mobility-driven compositional change. We quantify the population response to growth shocks directly in Section 5.3.2 and discuss the implications for interpreting the citizenship results in Section 5.3.4.

Two data challenges deserve mention. First, NUTS-2 regional definitions changed over our sample period, with regions being split or reallocated. Where possible, we re-merge split regions

food service; J: Information and communication; K: Financial and insurance; L: Real estate; M-N: Professional, scientific and technical; administrative and support service; O-Q: Public admin., defence, education, human health and social work; R-U: Arts, entertainment and recreation; other services; act. of households and extra-territorial. Note that Manufacturing is included not only in B,C,D,E category, but also separately. Here we omit data on just the manufacturing industry and include it bundled with Mining, Energy, Manufacturing and Water supply. Therefore, we have 10 non-overlapping industries.

[4] As a robustness check we also run our regressions constructing the Bartik instrument using the 15 countries from our regression, and produce virtually identical results.

to maintain a consistent panel; where changes are too complex to reconcile, we truncate the series at the point of redefinition. Second, the LFS classification of educational attainment shifted from ISCED 1997 to ISCED 2011 in 2014, creating a break in the treatment of short-cycle tertiary education and cross-country inconsistencies. We address this by grouping education into three categories: lower secondary, upper secondary, and tertiary education, the categories are robust to the reclassification.

Regional GDP and GVA data are used to construct the business cycle shock and the Bartik instrument, respectively. GVA is available by NACE industry classification, which we aggregate into 10 broad non-overlapping sectors. The panel is not perfectly balanced across years, as some countries begin reporting GVA data later than 2000; we account for this through the inclusion of year fixed effects. Finally, to reduce the impact of outliers, we trim 5% of the sample symmetrically by excluding observations below the 2.5th percentile and above the 97.5th percentile.

Descriptive evidence on reasons for inactivity. The LFS contains questions that provide useful context for interpreting our results.

Main reason for not wanting to work	% of responses
Retirement	47.49%
Education or training	23.47%
Own illness or disability	10.36%
Other family reasons	4.88%
Other personal Reasons	4.76%
Care responsibilities	4.70%
Other reasons	4.34%

Table 2: Main reason participants outside the labour force state they are no longer looking for work. Source: Labour Force Survey.

Table 2 shows that among respondents outside the labour force, the most common reason for not wanting to work is retirement (47%), followed by education or training (23%) and illness or disability (10%). Table 3 disaggregates reasons for leaving the last job by labour force status. For those who remain in the labour force (i.e., the unemployed), the dominant reasons are involuntary: fixed-term job ended (44%) and dismissal or business closure (38%). By contrast, for those outside the labour force, retirement accounts for 47% of exits. This sharp contrast, over 80% of those remaining in the labour force were ‘forced’ out of their jobs versus largely voluntary exits among the inactive, provides important context for understanding which margin of adjustment is cyclical versus structural.

Main reason for leaving last job or business	Outside the Labour Force	Inside the Labour Force
Dismissal or business closed	11.56%	37.61%
Fixed-term job ended	13.40%	43.51%
Care responsibilities	3.91%	1.14%
Other family reasons	0.13%	0.08%
Education or training	2.71%	1.25%
Own illness or disability	12.16%	2.57%
Retirement	47.29%	0.64%
Other Personal reasons	0.30%	0.36%
Other reasons	8.56%	12.84%

Table 3: Main reason participants in the LFS opt to leave their last job or business. Data is split into those who are in the labour force (i.e., unemployed) and those who are not in the labour force. Source: Labour Force Survey.

5 Results: Growth Shocks

5.1 Aggregate Responses

Figure 1 presents the impulse response functions (IRFs) of our three age-sex adjusted outcome variables to a temporary 1 percentage point decrease in GDP growth. The employment rate declines contemporaneously by about 0.1 percentage points, reaching a trough of -0.41 percentage points after three years, and recovers to pre-shock levels after approximately five years. The unemployment rate mirrors this pattern: it rises immediately by 0.17 percentage points, peaks at 0.5 percentage points after two years, and also recovers within five years. These magnitudes and dynamics are consistent with the well-documented persistence of unemployment in European labour markets (Caporale et al., 2021; Ball, 2014) and with the literature documenting higher unemployment persistence in Europe than in the U.S. (Alogoskoufis and Manning, 1988).

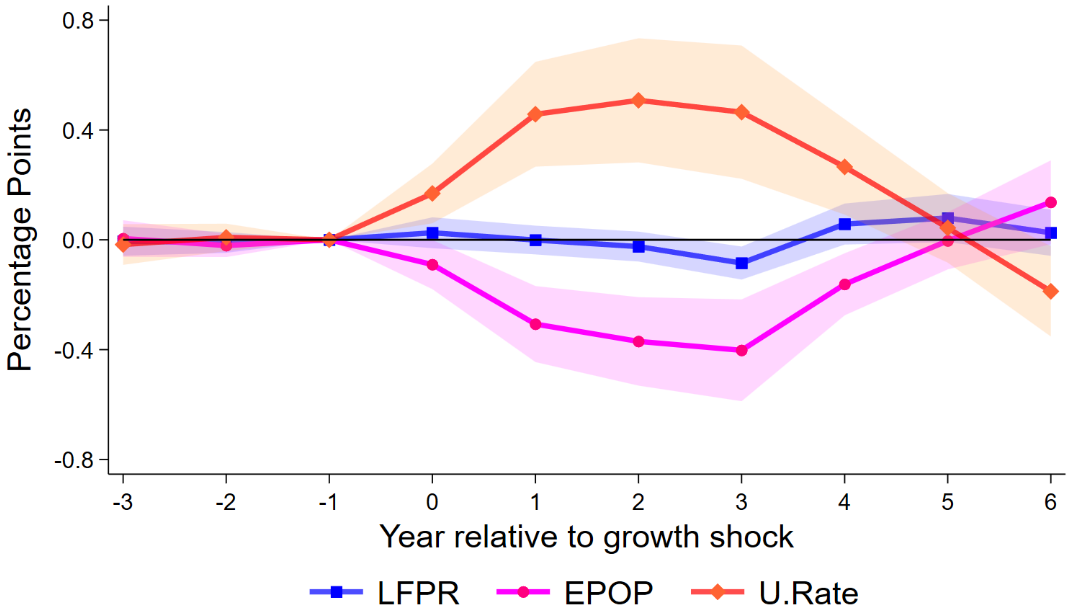

Figure 1: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange). Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for $k = 0$: 16.5.

The central finding is the insensitivity of the LFPR to growth shocks. With the exception of a marginally negative and significant reaction of -0.1 percentage points in year three, overall the LFPR does not respond significantly to a 1 percentage point output shock. While the resilience of the EU LFPR has been documented descriptively (Bodnár, 2018; Balleer et al., 2009), our paper provides the first causal estimate of this resilience using exogenous output shocks identified through a shift-share instrument. This result contrasts sharply with the U.S., where Cajner et al. (2025) find that the LFPR declines for four years and takes eight years to fully recover. Because we employ a methodology comparable to theirs, this divergence can be attributed to differences in labour market structures rather than differences in empirical approach. Several institutional features may account for this EU–U.S. divergence. Stronger employment protection legislation in the EU reduces involuntary separations, limiting the flow from employment into inactivity. Moreover, more generous social insurance, unemployment benefits, short-time work schemes and active labour market programmes help to sustain job search activity and hence labour force attachment during downturns.

We formally test whether the LFPR and unemployment rate follow the same time profile by running a pooled regression and testing whether the ratio $\beta^{(k)}_{\text{UR}}/\beta^{(k)}_{\text{LFPR}}$ is constant across horizons. A nonlinear Wald test strongly rejects this null ($\chi^2(7) = 33.75$, $p < 0.000$), confirming that the two outcome variables follow distinct time profiles.

In Table 4, we report the F-statistics for LFPR, the employment and unemployment rates across the horizons 0–6. The results signify that GVA growth provides a strong instrument for GDP growth at short and medium horizons, satisfying the relevance condition. We perceive a gradual decline in the F-statistics at longer horizons, reflecting the diminishing predictive power of contemporaneous industry shocks for longer-run outcomes, as additional sources of variation and adjustment dynamics become more important over time. Despite this decline, the instrument remains sufficiently strong across most horizons to support reliable IV estimation, although estimates at longer horizons should be interpreted with greater caution. Overall, GVA growth provides a credible source of exogenous variation for identifying GDP shocks.

Horizon	LFPR	EPOP	URATE
0	16.5	15.37	17.5
1	22.9	20.6	23.3
2	22.0	20.3	22.9
3	20.4	19.6	22.1
4	17.7	17.5	19.7
5	12.5	12.6	14.3
6	8.4	8.5	9.8

Table 4: Reported F-statistic for horizons 0–6 of IRF.

Historical context. Figure 2 provides descriptive context by plotting the age-sex adjusted LFPR, employment rate, and unemployment rate from 1990 to 2022. The LFPR and employment rate tracked each other closely until the Great Recession, when a sharp divergence emerged: the LFPR continued its gradual increase while the employment rate dropped sharply and the unemployment rate surged. This pattern, rising participation alongside falling employment, suggests that even during the largest economic shock in the sample, individuals in the EU continued to enter or remain in the labour force, either actively searching for work or maintaining their availability. This descriptive evidence is consistent with our econometric finding that the LFPR does not respond to cyclical shocks, and motivates the deeper investigation of which demographic groups are driving this resilience.

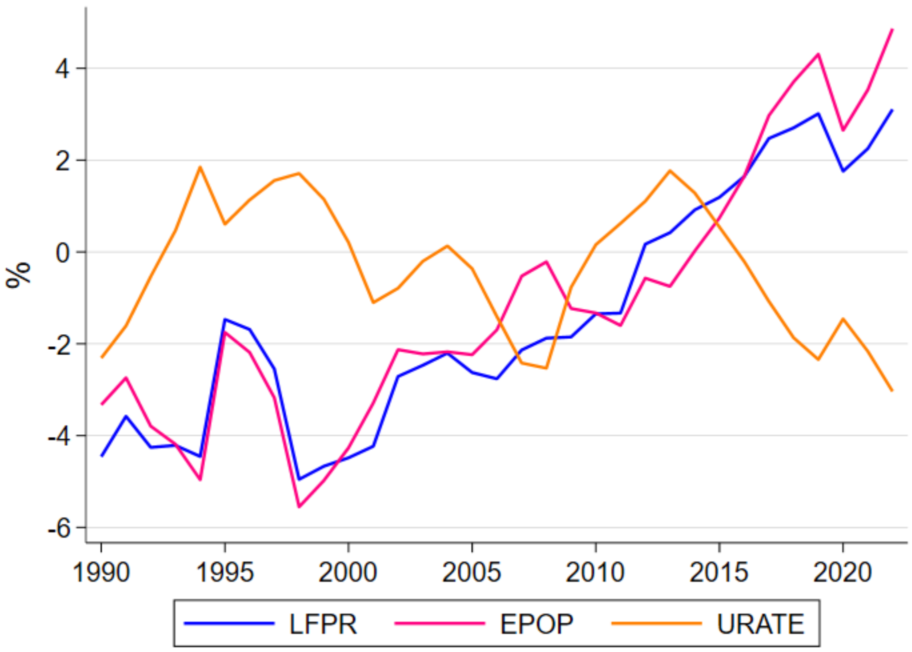

Figure 2: Age-adjusted LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (URATE, orange) 1990–2022. The vertical axis denotes the rates.

How important is the age-sex adjustment? Figure 3 compares the IRFs of the unadjusted LFPR and age-sex adjusted LFPR. Both specifications yield qualitatively similar results: the LFPR is insensitive to growth shocks regardless of the adjustment. This suggests that the resilience of the aggregate LFPR is not an artefact of compositional changes masking underlying cyclical movements. We investigate further whether controlling for other important factors such as education, alters our inference. Our results are shown in the Appendix (Figure 13). Since the LFS only provides data for education from 2006 onwards, we estimate the age-sex adjusted, and age-sex-education adjusted LFPRs, along with their fitted values from 2006–2020 for consistency and proper comparison. We notice that adjusting for education does not significantly alter our results, suggesting that age-sex adjustment by itself adequately captures changes in labour market composition.[5]

[5]We observe a slight significant fall in the age-sex and age-sex-education adjusted LFPR, because we use only 2006–2020 data, which reduces our sample.

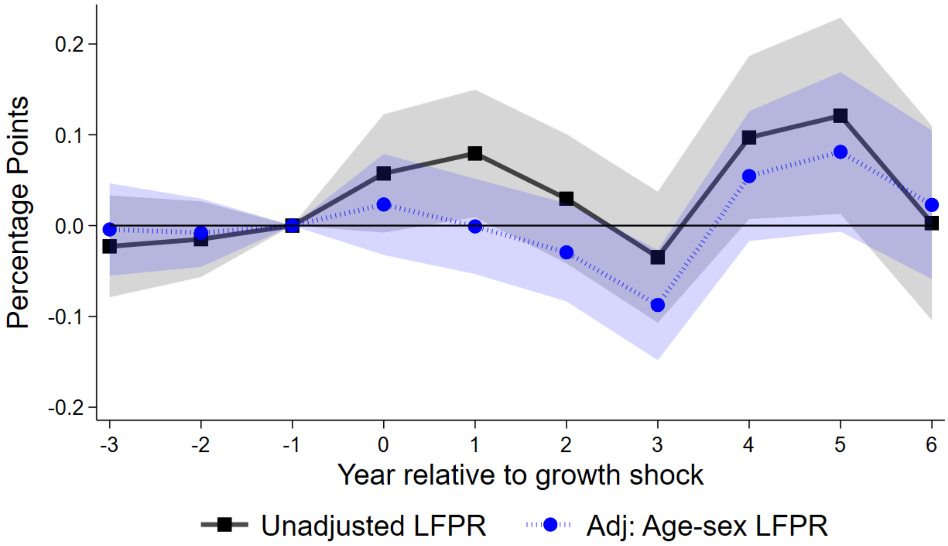

Figure 3: Estimated IRFs for unadjusted and age-sex adjusted LFPR. The bands around each line indicate a 95% confidence interval for the age-sex adjusted LFPR, representing standard errors clustered at the regional level. $k = 0$: 17.2 for unadjusted LFPR and 16.5 for age-sex adjusted LFPR.

5.2 Mechanisms: Labour Market Flows and Reasons for Inactivity

To shed light on the mechanisms underlying aggregate LFPR resilience, we exploit LFS questions on reasons for leaving the last job. Estimating Equation 1 with these reasons as outcome variables, we find that the share of respondents citing dismissal or business closure rises sharply and persistently (up to four years) after a negative growth shock (Figure 4). Other reasons, including fixed-term job ended, illness, and education, decrease in relative terms, indicating that the cyclical margin is dominated by involuntary separations rather than voluntary exits. Notably, retirement shows no significant cyclical response, suggesting that early retirement is not a primary margin of adjustment to growth shocks in our sample.

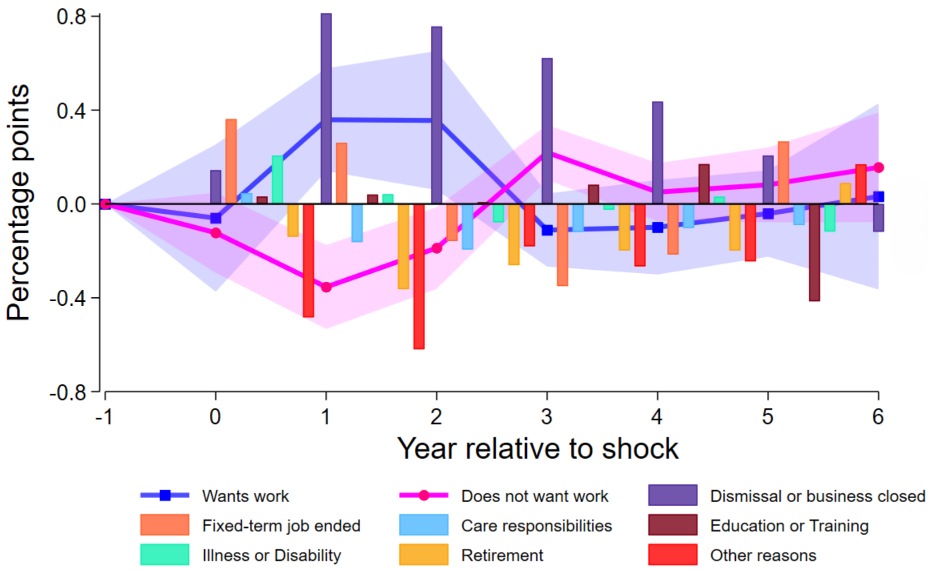

Figure 4: Blue line shows the estimated IRFs for the share of individuals who want to work even if they are not searching for employment. The pink line indicates the share of the population who does not want to work and is not looking for employment. The bars plot the coefficients from Equation 1, where our outcome variable is the share of the population stating the main reason for leaving their last job or business. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands indicate a 95% confidence interval, with standard errors clustered at the regional level.

We also examine the share of inactive individuals who express a desire to work. This indicator, which can be interpreted as a measure of the "shadow" labour force or labour market slack beyond the unemployment rate, increases after negative growth shocks, consistent with the discouraged worker effect operating on the extensive margin. The effect lasts for a few years before dissipating. This is relevant for monetary policy: it suggests that even when the LFPR appears stable, there may be hidden slack that the unemployment rate does not capture.

5.3 Heterogeneous Responses in Demographics

5.3.1 Gender

Figure 5 reveals a striking gender asymmetry. Men's LFPR declines significantly following the shock, reaching a trough of -0.15 percentage points after three years. Their employment rate falls, reaching a trough of -0.52 percentage points after three years. The unemployment rate rises, peaking two years after the shock. Both the employment rate and unemployment rate recover approximately five years after the initial shock. Women's LFPR, by contrast, shows no significant response at any horizon. Their employment rate and unemployment rate do respond, but with smaller magnitudes and slightly faster recovery, four years versus five for men.

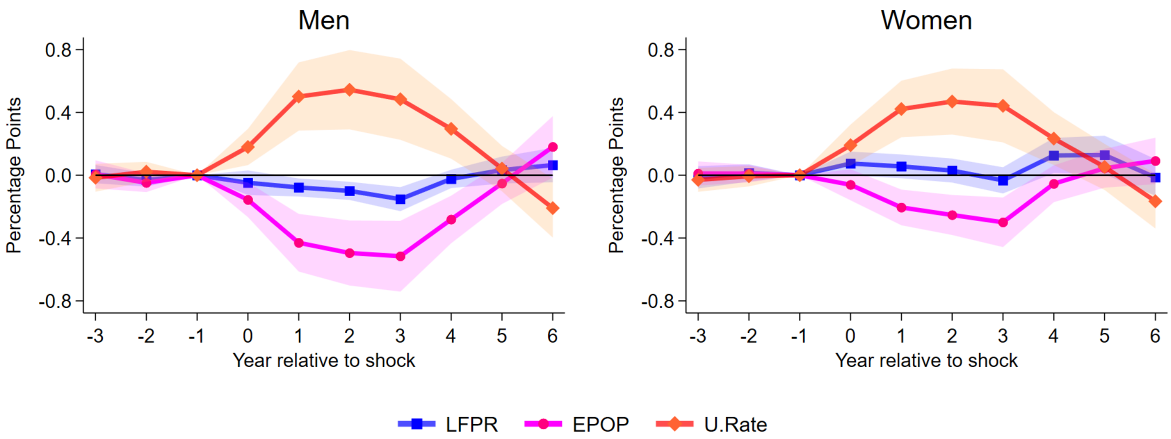

Figure 5: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) by **gender**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands indicate a 95% confidence interval, with standard errors clustered at the regional level. F-statistic for k = 0: 13.8 for men and 13.9 for women.

Three interrelated factors help explain this pattern. First, sectoral segregation: men are overrepresented in cyclically sensitive industries, construction and manufacturing, while women are concentrated in more stable sectors such as education, healthcare, and public administration. Verdugo and Allègre (2020) reach a similar conclusion using occupation-based Bartik instruments for 10 EU countries during the Great Recession, finding that female LFPR increased more in regions where male-dominated sectors contracted most sharply. Our results confirm their gender finding using a different identification strategy (industry-based output shocks rather than occupation-based employment shocks), a broader sample (15 countries, 2000–2020), and a wider set of outcome variables. Crucially, because we estimate all demographic subgroups within the same framework, we can show that the gender asymmetry interacts with the age, education, and industry dimensions, something the existing literature has not done. Second, the gender wage gap and the prevalence of part-time and flexible work among women reduce firms' incentive to dismiss female employees during downturns (Gomes, 2024; Périvier, 2014). Third, the added worker effect: as household income falls due to male job loss, married women enter or remain in the labour force to compensate (Engemann and Wall, 2009).

The strength of the added worker effect varies substantially across European welfare state regimes. Riedl and Schoiswohl (2015) document a strong effect during the Great Recession, with women constituting the majority of added workers. However, this effect appears concentrated in Mediterranean countries that provide lower welfare benefits (Bredtmann et al., 2017; Prieto-Rodriguez and Rodriguez-Gutierrez, 2003), while Continental European and Nordic countries show little or no effect (Hardoy and Schøne, 2014; Cammeraat et al., 2023). This suggests that the net effect on aggregate LFPR depends on the welfare state configuration, with state-provided insurance and spousal labour supply acting as substitute mechanisms for smoothing household income.

5.3.2 Age

Figure 6 shows the responses by age group: 15–24, 25–54, and 55–74. The young are hit hardest across all three outcome variables. Their unemployment rate peaks at 0.6 percentage points three years after the shock, and their employment rate trough reaches −0.43 percentage points after three years. The prime-age group's unemployment rate peaks earlier (after two years, at 0.5 percentage points). The older generation is least affected, with an employment rate trough of −0.3 percentage points after two years.

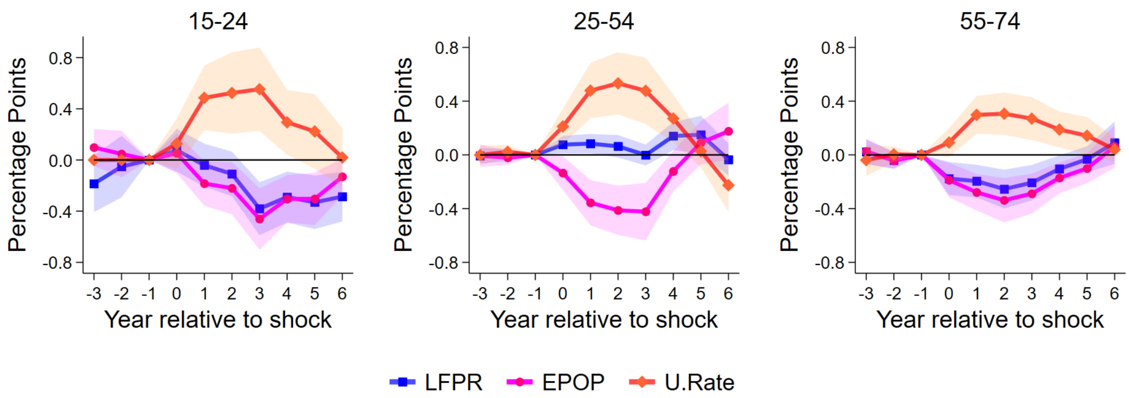

Figure 6: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) **by age**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for $k = 0$: 33.2 for young, 16.8 for prime-aged and 14.8 for older workers.

Most importantly, the LFPR responses differ qualitatively across age groups. The prime-age LFPR is unaffected, consistent with strong labour force attachment driven by family responsibilities and mortgage obligations. The young exhibit a significant and highly persistent LFPR decline, with effects lasting up to eight years. Given the inclusion of region-specific time trends, this persistence is identified as a shortfall relative to each region's own pre-shock secular path (see Section 3.1). Accordingly, we interpret it as medium-run scarring at horizons relevant for human capital formation and labour market entry, rather than as a permanent reduction in long-run equilibrium participation. The older group also shows a significant LFPR decline, consistent with a discouraged-worker effect among older workers when conditions deteriorate; we return to the role of early retirement below.

The persistent LFPR decline among the young is particularly concerning and connects to the broader literature on labour market hysteresis. Since Blanchard and Summers (1986), it has been recognised that demand shocks can have permanent effects on the equilibrium unemployment rate, and, by extension, on labour force participation, if cyclical dislocations erode human capital, weaken labour market attachment, or alter the matching process. Our findings suggest that this hysteresis channel operates with particular force among the young, consistent with Bell and Blanchflower (2011), who document that youth unemployment is negatively correlated with participation, reflecting a strong discouraged worker effect. The young are more likely to lose their jobs, due to lower skills, less experience, and seniority-weighted redundancy costs, and more likely to leave the labour force entirely, often retreating into education (Dellas and Koubi, 2003). Yagan (2019) provides direct evidence of employment hysteresis from the Great Recession in the U.S., finding that exposure to a 1 percentage point larger local unemployment shock reduced employment rates six years later by 0.3 percentage points, even after unemployment rates had signalled recovery. Oreopoulos et al. (2012) document that entering the labour market during a recession imposes persistent wage penalties lasting a decade, with effects concentrated among lower-skilled graduates.

Older workers also experience a significant fall in their LFPR, with a contemporaneous decline of 0.17 percentage points that reaches its trough of -0.25 percentage points two years after the shock. This likely reflects a discouraged worker effect: older workers who lose their jobs face age discrimination in hiring, fewer suitable opportunities, and skill mismatch, leading them to withdraw from job search (Carlsson and Eriksson, 2019; OECD, 2019). Whilst Axelrad et al. (2017) find that during the Great Recession older workers were more likely to remain unemployed than to retire, our results suggest that some nonetheless exit the labour force entirely. Early retirement may also play a role, as firms use workforce reductions that disproportionately affect older employees (Johnson, 2012; Coile et al., 2012). The results above use age-sex adjusted outcome variables; unadjusted results are shown in the Appendix, Figure 14. For young and prime-aged workers, the unadjusted results remain comparable, highlighting that we mainly face differences *across* these age groups rather than *within* them. The same, however, does not hold for older workers. Interestingly, without the age-sex adjustment the LFPR ceases to show a significant fall, affirming the relevance of this adjustment. Structural changes within the older cohort, including an ageing population and older workers opting to work for longer, whether due to financial obligations or improved health, conceal their true cyclical responses to growth shocks, which we isolate through our framework.

Crucially, the distinction between demand shocks and other business cycle disturbances matters here. Our Bartik instrument isolates exogenous demand shocks at the regional level, and it is precisely these demand-driven downturns that the hysteresis literature identifies as most damaging for long-run labour market outcomes. When demand contracts, firms shed young workers first; the resulting unemployment spells erode human capital, weaken professional networks, and may induce permanent withdrawal from the labour force (Cockx and Ghirelli, 2016; Acabbi et al., 2024). In the Appendix, we observe a structural break in youth LFPR commencing during the Great Recession depicted in Figure 15, with the pre-crisis level never fully recovering. This is consistent with the hysteresis mechanism suggested in the literature: a sufficiently large demand shock can permanently alter a cohort's relationship with the labour market, with long-run consequences for aggregate labour supply, potential output, and, through its effect on the NAIRU, the conduct of monetary policy (Galí, 2020; Ball et al., 2022).

Moving into inactivity is one option for people. Another option is to leave the region they are living, in search of better jobs elsewhere. We can test this channel by changing the left-hand-side variable from LFPR to population per region and analyse the age groups' different reactions after a shock. We do find significant migration responses to growth shocks. Figure 7 shows the effect of the shock on log population by single-year age group. There is a contemporaneous and persistent outflow of individuals aged 25–35 from economically distressed regions, with effects lasting five years or more. Interestingly, those in their mid-sixties and over also tend to leave affected regions, possibly due to reduced economic support in older age. These findings underscore the importance of the age-sex adjustment, as migration-induced compositional changes in the regional population would otherwise contaminate estimates of cyclical participation responses (Blanchard and Katz, 1992; Barslund et al., 2014; Cavalleri et al., 2021).

The age-sex adjustment cannot, however, absorb selection *within* demographic cells. If the individuals leaving distressed regions have systematically higher participation rates than observationally similar individuals who stay (plausible, given that mobility is itself an investment more attractive to those attached to the labour market), the LFPR of the remaining population could appear stable even as genuine withdrawal from activity occurs among movers. Part of the aggregate resilience we document may therefore reflect selective out-migration rather than uniform stability among stayers. Our annual, repeated cross-section design does not allow us to track individuals across regions or borders, so we cannot quantify this channel directly; we flag it as a caveat to interpretation and return to it in the context of citizenship in Section 5.3.4.

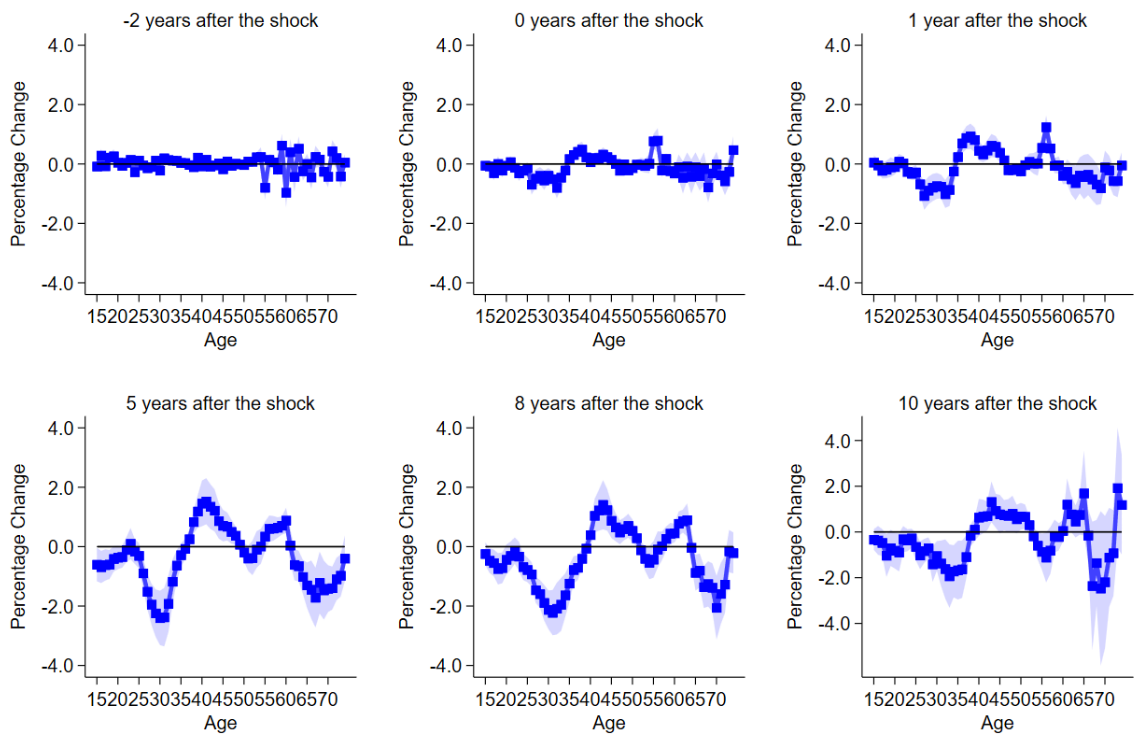

Figure 7: Estimated IRFs for **population change**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, with standard errors clustered at the regional level.

5.3.3 Education

Among prime-age workers, those with at most upper secondary education experience significantly larger employment rate and unemployment rate responses than tertiary-educated workers (Figure 8). The employment rate trough for the less-educated group reaches -0.37 percentage points after three years, and their unemployment rate peaks at 0.7 percentage points after two years. Tertiary-educated workers show smaller and less persistent responses. These findings reflect the multifaceted vulnerability of less-educated workers: they are more likely to hold temporary contracts, work in volatile industries, and possess skills that are easier to substitute (Cedefop, 2017; Hoynes et al., 2012). It is primarily low-educated men who are most at risk (Berghammer and Adserà, 2022).

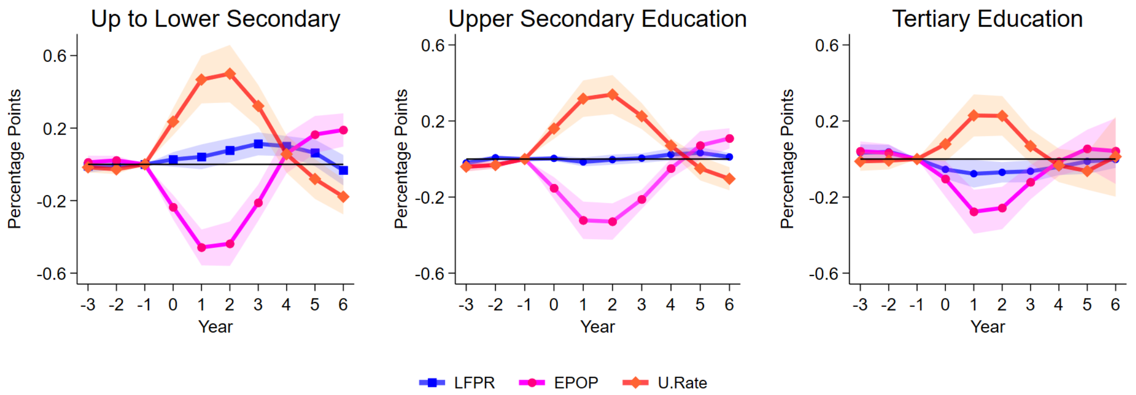

Figure 8: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) - **by education**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for k=0: 33.13 for less than upper secondary education, 23.10 for upper secondary and 24.83 for tertiary education.

5.3.4 Citizenship

Non-EU citizens are the most affected group (Figure 9). Their unemployment rate peaks at 0.96 percentage points two years after the shock, and their employment rate falls to -0.77 percentage points. EU citizens show intermediate responses, and natives the smallest. The LFPR is not significantly affected for any citizenship group. These results align with the "ethnic penalty" literature (Reyneri and Fullin, 2011; Illing and Koch, 2021; Bratsberg et al., 2016) that suggests there remains a difference in immigrants' and natives' labour market outcomes when socio-demographic characteristics are controlled for. The picture is heterogeneous across countries: nations importing primarily highly skilled migrants saw non-natives less affected (OECD, 2011; Canagarajah, 2011), while countries experiencing surges in less-skilled migration saw the opposite (Zorlu and Hartog, 2008).

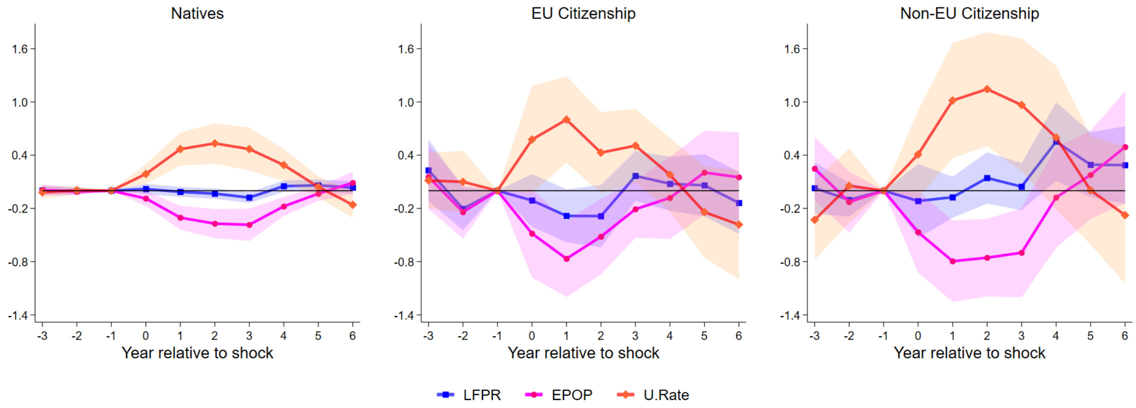

Figure 9: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) - **by citizenship**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for $k = 0$: 16.7 for natives, 17.9 for EU citizens, and 11.8 for non-EU citizens.

The distinction drawn in Section 4 between citizenship as a demographic attribute and migration as a population flow matters for interpreting these results: the non-response of the LFPR across citizenship groups may partly reflect return migration rather than stable participation among a fixed population. When non-native workers lose their jobs during a downturn, some may leave the country rather than transition into inactivity within it, particularly non-EU citizens whose residence status may be conditional on employment (Koehler et al., 2011). Because such individuals exit both the numerator and denominator of the LFPR simultaneously, their departure leaves the participation rate mechanically unchanged even if a genuine withdrawal from economic activity has occurred. Bratsberg et al. (2016) document that immigrants from low-income countries face a high probability of permanent labour market exit during economic downturns, with emigration constituting one such exit route. The migration responses documented in Figure 7, which show persistent outflows of working-age individuals from economically distressed regions, are consistent with this interpretation (Barslund et al., 2014; Cavalleri et al., 2021). In the Appendix, Figure 16, we split the data into Natives vs Non-natives, grouping EU and non-EU citizens together. By doing so, we add more data weight to the latter group given the majority of people in the LFS identify as native. We see our results, and hence conclusions, remain unchanged.

5.4 Industry

We group industries into primary (agriculture and mining), secondary (manufacturing, utilities, and construction) sectors versus the tertiary (services) sector[6] to maintain statistical power given the shorter sectoral data (2008–2020). Figure 10 shows that the primary and secondary sectors exhibit significantly larger responses. Notably, there is a contemporaneous fall in the LFPR of −0.25 percentage points, with effects lasting four years. The tertiary sector shows more moderate responses and no LFPR effect.

These industry-level results provide the structural underpinning for the gender asymmetry. The "mancession" pattern, men disproportionately affected due to their concentration in construction and manufacturing (Wall, 2009; Arpaia and Curci, 2010), is directly reflected in the primary/secondary sector LFPR decline. The relative insulation of the tertiary sector, where female employment is concentrated, helps explain why women's LFPR remains stable (Eurofound, 2011; Verdugo and Allègre, 2020).

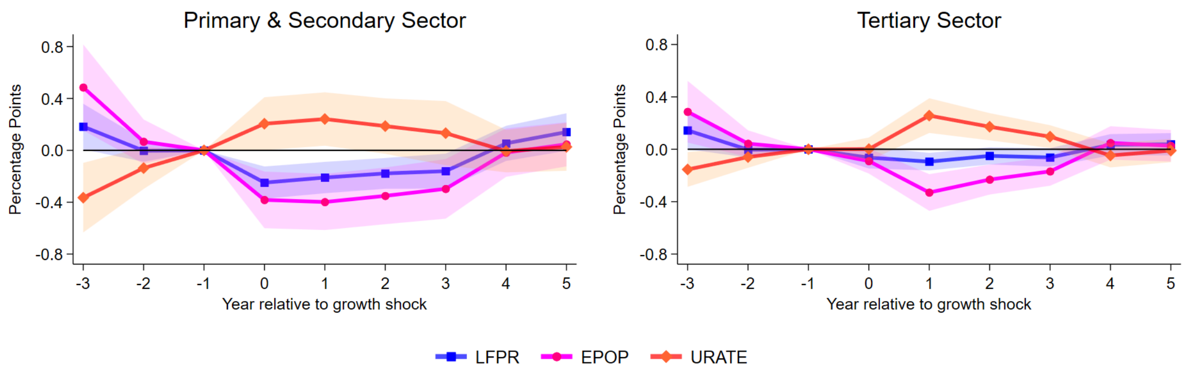

Figure 10: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) - **by sector**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for $k = 0$: 28.2 for primary & secondary sectors and 21.6 for tertiary sector.

[6]The LFS categorises industries in accordance with NACE. The data up until 2020, which is the sample we use, follows NACE Rev.2. The NACE categories were revised and implemented in October 2022 (NACE Rev.2.1). Primary Sector includes: Agriculture, forestry and fishing and Mining and quarrying. The Secondary Sector includes: Manufacturing, Electricity, Gas, Steam and Air-Conditioning Supply, Water supply and Construction. The Tertiary Sector includes: Wholesale and Retail Trade, Repair of Motor vehicles and Motorcycles, Accommodation and Food Service Activities, Transportation and Storage, Information and Communication, Financial and Insurance Activities, Real Estate Activities, Professional, Scientific and Technical Activities, Administrative and Support Service Activities, Public Administration and Defence, Compulsory Social Security, Education, Human health and Social Work Activities, Arts, Entertainment and Recreation, Other Service Activities, Activities of Households as Employers, Undifferentiated goods and services - producing activities of households for own use and Activities of Extraterritorial Organisations and Bodies.

5.5 Taking Stock: The Value of a Unified Framework

Many of the individual patterns documented above have antecedents in separate literatures. Table 5 summarises, for each demographic dimension, what was previously known and what our unified framework adds. Three contributions cut across all dimensions. First, all responses are identified from the same exogenous growth shocks, in the same specification, over the same sample—so differences across groups reflect genuine heterogeneity rather than differences in methods, samples, or shock definitions across studies. Second, the framework permits an explicit aggregation exercise: the muted response of the aggregate LFPR can be traced to offsetting subgroup dynamics (stable women and prime-age workers against responsive men, youth, and older workers), a decomposition unavailable when each margin is studied in isolation. Third, interactions across dimensions become visible, most notably that the gender asymmetry operates through sectoral exposure, and that the education gradient is concentrated among men.

Dimension	Existing evidence	This paper
Gender	Female LFPR rose in EU regions where male-dominated sectors contracted (Verdugo and Allègre, 2020); added worker effect varies across welfare regimes (Bredtmann et al., 2017; Hardoy and Schøne, 2014)	Causal confirmation over 2000–2020; gender asymmetry shown to operate through sectoral exposure (Section 5.4), estimated jointly with all other margins
Age	Youth unemployment negatively correlated with participation (Bell and Blanchflower, 2011); persistent wage scars from recession entry (Oreopoulos et al., 2012); U.S. employment hysteresis (Yagan, 2019)	First causal estimates of youth LFPR persistence in the EU, identified as deviation from region-specific trends; older workers' response only visible after age-sex adjustment
Education	Less-educated workers bear disproportionate recession costs in the U.S. (Hoynes et al., 2012)	EU evidence within the same framework; effects concentrated on employment and unemployment margins rather than participation; gradient strongest among men
Citizenship	"Ethnic penalty" amplified in recessions (Reyneri and Fullin, 2011; Bratsberg et al., 2016)	Largest employment and unemployment responses among non-EU citizens; LFPR resilience interpretable through return migration exiting numerator and denominator simultaneously (Section 5.3.4)
Industry	"Mancession" pattern of male overrepresentation in cyclical sectors (Wall, 2009; Arpaia and Curci, 2010)	Provides the structural underpinning of the gender asymmetry: significant LFPR fall in primary/secondary sectors, none in services

Table 5: Heterogeneous LFPR cyclicality: existing evidence and the contribution of a unified causal framework.

6 Robustness and Instrument Validity

6.1 Productivity Decomposition

Following Cajner et al. (2025), we decompose the output shock into its employment and productivity components. This serves two purposes: it validates that our instrument captures genuine output variation rather than merely proxying for employment changes, and it quantifies how much of the growth shock operates through the productivity channel, a margin that employment-based instruments would miss by construction.

Figure 11 plots the response of GDP and productivity (real GVA per hour worked) in both growth rates and levels. Productivity accounts for roughly half of the adverse growth shock. Productivity growth returns quickly (within one year), but the level takes approximately four years to recover, consistent with medium-term productivity effects of European recessions (Mattsson et al., 2020; Furceri et al., 2021). This also underscores the advantage of our growth-shock approach: an employment-based Bartik instrument would miss the productivity channel entirely (Cajner et al., 2025).

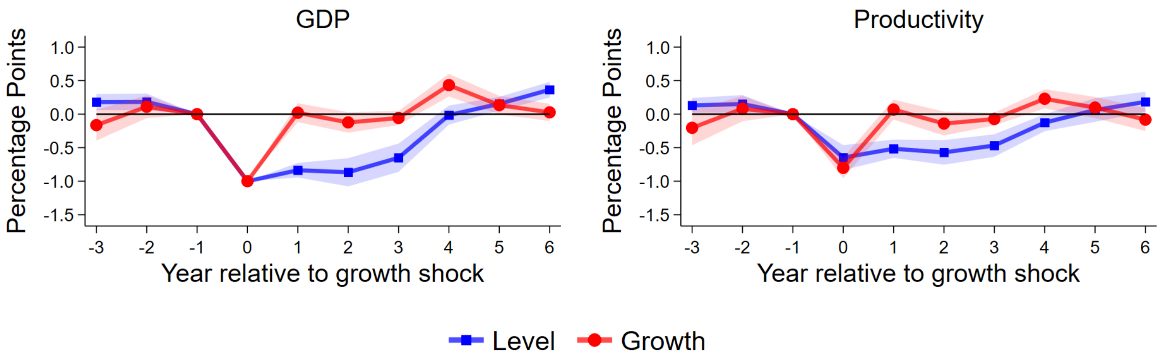

Figure 11: Productivity decomposition: response of GDP and productivity (measured as real GVA per hour worked) in both growth rates and levels.

6.2 Lead-Lag Exogeneity

We test the lead-lag exogeneity condition by estimating Equation 1 with the Bartik instrument as the outcome variable. Figure 12 confirms that the instrument responds only at horizon zero and shows no significant correlation at any lead or lag, validating the exogeneity assumption.

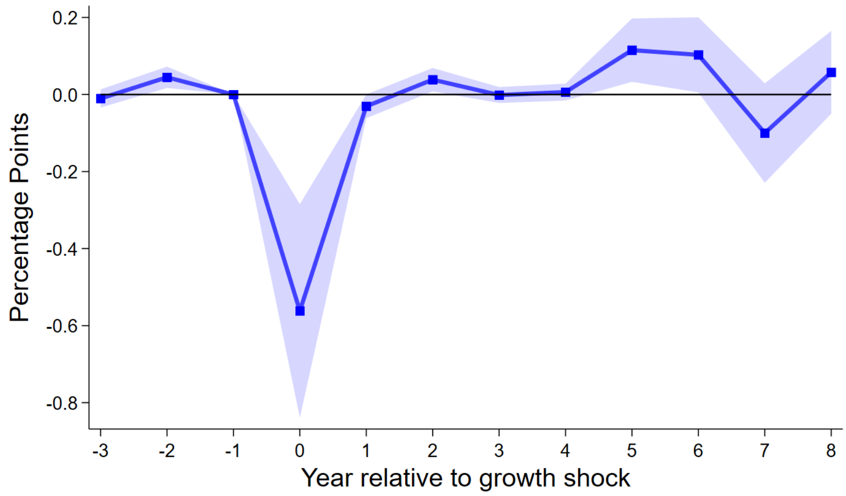

Figure 12: Estimating Equation 1 where the Bartik instrument is the outcome variable. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level.

6.3 Decomposition of Shift-Share Instrument

Following Goldsmith-Pinkham et al. (2020), we decompose the variation in the Bartik instrument into industry-specific Rotemberg weights to assess which industry-year observations are most influential in our estimates.

$$\hat{\alpha}_{kt} = \frac{g_{kt}\zeta'_{kt}\Delta GDP^{\perp}_{t,t-1}}{\sum_{k'}\sum_{t'} g_{k't'}\zeta'_{k't'}\Delta GDP^{\perp}_{t,t-1}}, \quad \hat{\beta}_{kt} = \frac{\zeta'_{kt}\Delta EPOP^{\perp}_{t+k,t-1}}{\zeta'_{kt}\Delta GDP^{\perp}_{t,t-1}}, \quad \hat{\beta} = \sum_k \sum_t \hat{\alpha}_{kt}\hat{\beta}_{kt} \tag{4}$$

where $\Delta GDP^{\perp}_{t,t-1}$ and $\Delta EPOP^{\perp}_{t,t-1}$ are the residualised GDP growth rate and EPOP respectively, after controlling for regional and year fixed effects, and ζ'_{kt} is the industry share for industry k. Given that the EU's LFPR exhibits strong resilience, we focus on the EPOP. In particular, we examine horizon 3, where the EPOP reaches its trough.

Our results are shown in Table 6, where we report the top eleven Rotemberg weights, $\hat{\alpha}_{kt}$. We observe that *Industry (except construction)* in 2009 carries a particularly large weight (1.03), reflecting the sharp contraction during the Great Recession. By then, *Industry (except construction)* had dropped by approximately 12%, and accounted for 17% of total GVA, highlighting the magnitude of this shock. Given substantial cross-regional variation in exposure to industrial activity, this generates strong covariance with regional outcomes and explains the prominence of this observation in the Rotemberg decomposition. Importantly, however, this high-weight observation does not drive our results.

Excluding this observation, we find that the subsequent top 10 weights contribute 71% of total weight in our shift-share instrument. This delivers a weighted $\hat{\beta}_{kt}$ of 0.49, which is broadly consistent with our baseline IV estimate of 0.40. The dominant industries are *Industry (except construction)*, *Agriculture, Forestry & Fishing*, *Construction*, and *Wholesale & Retail Trade*. We further assess robustness by excluding the two largest weights and focusing on the ten observations commencing from the third largest $\hat{\alpha}_{kt}$. We show the results in the appendix, Table 7. These contribute 52% of the total weight in our instrument and deliver a weighted $\hat{\beta}_{kt}$ of 0.42, which is virtually identical to our baseline IV estimate. Finally, in the robustness check, aggregating all remaining industry-year cells (excluding the largest observation) yields a weighted $\hat{\beta}_{kt}$ of 0.48. Taken together, the weights are distributed across multiple industries and years, confirming that our results cannot be traced to a single shock.

Industry	**Year**	α_{kt}	β_{kt}
$\boxed{\text{Industry (except construction)}}$	$\boxed{2009}$	$\boxed{1.03}$	$\boxed{-0.29}$
Industry (except construction)	2010	0.22	0.16
Agriculture, Forestry & Fishing	2009	0.12	0.10
Industry (except construction)	2011	0.06	0.06
Construction	2006	0.06	-0.06
Professional, Scientific & Technical Activities	2011	0.05	0.04
Agriculture, forestry & fishing	2017	0.05	0.01
Construction	2010	0.05	0.03
Wholesale & Retail Trade	2019	0.05	0.03
Wholesale & Retail Trade	2012	0.03	-0.01
Industry (except construction)	2003	0.03	-0.02

Table 6: Rotemberg weights in GDP Shift-Share Instrument used in our main estimates.

6.4 Additional Robustness Checks

We assess the sensitivity of our findings to several alternative specifications. First, we vary the lag structure of the industry shares used to construct the Bartik instrument. Our baseline uses a three-year moving average lagged five years. Adopting a three-year ($\omega_{q,r,t-3}$) or seven-year ($\omega_{q,r,t-7}$) lag instead of $\omega_{q,r,t-5}$ in Equation (2) yields virtually identical results, confirming that the choice of lag does not drive our conclusions. We show our results for $\omega_{q,r,t-3}$ in Figure 17 in the Appendix.[7]

Second, we assess the importance of region-specific time trends. As discussed in Section 3.1, a potential concern is that linear trends could absorb part of a slow-moving cyclical response, mechanically dampening the persistence we estimate (particularly relevant for the long-lived youth LFPR response). The evidence points in the opposite direction. Omitting the trends does not amplify the persistent effects; instead, it leads to a loss of significance and altered point estimates (Figure 18 in the Appendix), consistent with the trends absorbing secular regional divergence that would otherwise add noise rather than masking cyclical persistence. Adding quadratic region-specific trends, by contrast, leaves the results materially unchanged, indicating that the linear specification is sufficiently flexible. Together, these results suggest that the persistent responses we document (and the youth scarring in particular) are deviations from regional trend paths rather than artefacts of the trend specification.

Third, we examine whether the inclusion of the year 2020, during which the unprecedented COVID-19 pandemic struck, biases our estimates. Re-estimating our baseline specification excluding 2020 (Figure 19 in the Appendix) confirms that our results are robust to this exclusion.

Finally, we examine the sensitivity of inference. Beyond the baseline regional clustering motivated in Section 3.1, we report three alternative procedures.

First, since NUTS-2 regions are nested within countries with shared institutions and, for euro-area members, a common monetary policy stance, we re-cluster at the country level. Estimates retain their sign and approximate magnitude, with somewhat wider confidence bands reflecting the smaller number of clusters (Figure 20 in the Appendix).

Second, we cluster two-way by region and year, allowing for arbitrary correlation of residuals among regions in the same calendar year beyond what is absorbed by year fixed effects. Results for the aggregate specification are reported in Figure 22, with the demographic subsamples we examine under this procedure (gender, age) in Figures 24 and 26. Confidence intervals widen substantially relative to the baseline, reducing statistical significance at longer horizons; however, the key effects remain significant during the periods of strongest response.

We note that under both coarser procedures the first-stage F-statistic falls relative to the baseline value of 16.5, to 4.5 under country clustering and 4 under two-way clustering. This reflects the variance estimator rather than a genuine weakening of the instrument: the first-stage coefficient is unchanged, and coarser clustering produces a more conservative (larger) variance

[7] Results using $\omega_{q,r,t-7}$ are not shown for brevity, but are available from the authors upon request.

estimate that mechanically lowers the F, with the cluster-robust variance itself imprecise when there are as few as fifteen country clusters. It should therefore not be read as evidence of weak identification. The exposure-robust procedure of Adão et al. (2019) applied next is, moreover, the appropriate inference method for a shift-share design: its validity rests on the presence of many industry-level shocks rather than on a large number of geographic clusters, which is why we treat it as our preferred robustness check.

Third, and most importantly given our shift-share design, we apply the exposure-robust inference procedure of Adão et al. (2019), designed specifically to address dependence induced by common industry-level shocks weighted by region-specific shares. Aggregate results appear in Figure 21; demographic subsamples in Figures 23 (gender), 25 (age), 27 (education), 28 (citizenship), and 29 (industry).

Across all three procedures the qualitative pattern is preserved: the aggregate LFPR continues to show no significant response, the gender asymmetry and the persistent youth response remain identifiable, and the cancellation pattern across demographic subsamples carries through. The resilience of our results to the AKM correction is consistent with the Rotemberg decomposition (Section 6.3), which shows that identifying variation is spread across many industry-year cells rather than concentrated in a small number of shocks, precisely the condition under which the gap between conventional and AKM standard errors is smallest.

7 Conclusions

This paper provides the first causal estimate of the cyclical response of labour force participation to growth shocks across EU regions, using a methodology directly comparable to the U.S. evidence of Cajner et al. (2025). Using a shift-share instrument with a leave-one-country-out construction and local projections on EU Labour Force Survey microdata covering 15 countries and 114 regions over 2000–2020, we find that the aggregate EU LFPR shows no significant response to growth shocks at any horizon, whereas the employment rate and unemployment rate respond significantly and persistently, with effects lasting four to five years. While this resilience has been documented descriptively in earlier work (Bodnár, 2018; Balleer et al., 2009), our causal design establishes it as more than a descriptive regularity.

This aggregate resilience, however, masks pronounced heterogeneity, and it is in the demographic decomposition that the paper makes its most distinctive contribution relative to the existing European literature. While Verdugo and Allègre (2020) document the gender asymmetry in participation responses during the Great Recession, and Bell and Blanchflower (2011) and others highlight youth vulnerability, no study has examined all major demographic dimensions (gender, age, education, citizenship, industry) within a single causal framework. This integrated approach reveals that the aggregate LFPR resilience is not the result of uniform insensitivity across groups, but rather the product of offsetting dynamics: men's LFPR declines significantly while women's remains stable, driven by sectoral segregation and the added worker effect. Young workers exhibit the most persistent LFPR response (up to eight years), connecting to the hysteresis literature and suggesting potential scarring of labour supply. Less educated and non-native workers bear disproportionate employment costs, the latter partly through emigration. These subgroup findings are individually consistent with specialised studies, but demonstrating their coexistence and approximate cancellation in the aggregate is new.

Beyond this compositional offsetting, the stability reflects the margin along which European labour markets absorb shocks. In our estimates a negative shock is absorbed almost entirely through employment and unemployment, both of which respond for four to five years, while the pool of participants is left largely intact. Workers who lose their jobs tend to remain attached to the labour force as unemployed rather than leaving it, so that the discouraged worker channel that operates strongly in the U.S. is comparatively muted here. Our evidence on reasons for inactivity is consistent with this reading: the cyclical margin is dominated by involuntary separations, early retirement shows no cyclical response, and the rise in the "shadow" labour force is modest and short lived. Stronger employment protection, more generous unemployment insurance, and the widespread use of short-time work schemes plausibly sustain this attachment by keeping job losers within the measured labour force, and the more limited role of geographic mobility as an adjustment channel in Europe (Decressin and Fatás, 1995) works in the same direction. Because the net response also depends on how far state insurance substitutes for spousal labour supply (Bredtmann et al., 2017; Hardoy and Schøne, 2014), this institutional configuration helps explain why cyclicality varies across member states.

These findings bear on several policy debates. For monetary policy, the stability of aggregate participation implies that the unemployment rate is a comparatively reliable and timely gauge of cyclical slack in the EU, though the rise in the shadow labour force and the persistent, potentially scarring response of youth participation (Blanchard and Summers, 1986; Yagan, 2019) caution that some slack, and some erosion of future labour supply, may escape it. The same heterogeneity shapes fiscal and social policy, since adjustment concentrated among the young, the less educated, and secondary earners carries consequences for the tax base, for social expenditure, and for the design of targeted support during downturns.

We acknowledge several limitations. The rotating panel design precludes tracking individual transitions; matched employer-employee data could illuminate underlying flows. Regional fixed effects and time trends control for heterogeneity across regions but may also absorb some longer term effects of growth shocks, though our robustness checks indicate this does not drive the persistent estimates.

Several avenues for future research emerge. The role of specific labour market institutions in shaping cross-country LFPR cyclicality deserves systematic investigation. Extending the analysis beyond 2020 would test whether resilience patterns persist under qualitatively different shocks. Examining the interaction between older workers' rising participation and cyclicality would be valuable given the EU's demographic trajectory (Maestas et al., 2016). Finally, understanding the distributional consequences of heterogeneous participation responses would further inform social protection design.

References

Acabbi, E. D., Alati, A., and Mazzone, L. (2024). Human capital ladders, cyclical sorting, and hysteresis. Staff Working Paper 1077, Bank of England.

Adão, R., Kolesár, M., and Morales, E. (2019). Shift-share designs: Theory and inference. *The Quarterly Journal of Economics*, 134(4):1949–2010.

Alogoskoufis, G. S. and Manning, A. (1988). On the persistence of unemployment. *Economic Policy*, 3(7):427–469.

Arce, O. and Sondermann, D. (2026). Low unemployment, plenty of labour: what does it imply for wage pressures? Ecb blog, European Central Bank.

Arpaia, A. and Curci, N. (2010). Eu labour market behaviour during the great recession. Mpra paper, University Library of Munich, Germany.

Autor, D. H., Dorn, D., and Hanson, G. H. (2013). The china syndrome: Local labor market effects of import competition in the united states. *American Economic Review*, 103(6):2121–2168.

Axelrad, H., Sabbath, E., and Hawkins, S. (2017). The 2008-2009 great recession and employment outcomes among older workers. *Eur J Ageing*, 3;15(1):35–45.

Ball, L., Leigh, D., and Mishra, P. (2022). Understanding U.S. inflation during the COVID era. *Brookings Papers on Economic Activity*, pages 1–54.

Ball, L. M. (2014). Long-Term Damage from the Great Recession in OECD Countries. NBER Working Papers 20185, National Bureau of Economic Research, Inc.

Balleer, A., Gómez-Salvador, R., and Turunen, J. (2009). Labour force participation in the euro area: a cohort based analysis. Working Paper Series 1049, European Central Bank.

Barslund, M., Busse, M., Vargas-Silva, C., Kaczmarczyk, P., Baas, T., Peinado, M., Jimeno, J., Lacuesta, A., Galgóczi, B., and Leschke, J. (2014). Labour mobility in the EU: Dynamics, patterns and policies. *Intereconomics: Review of European Economic Policy*, 49(2):116–158.

Bartik, T. J. (1991). *Who Benefits from State and Local Economic Development Policies?* W.E. Upjohn Institute for Employment Research, Kalamazoo, MI.

Béland, L.-P., Brodeur, A., and Wright, T. (2020). The Short-Term Economic Consequences of COVID-19: Exposure to Disease, Remote Work and Government Response. GLO Discussion Paper Series 524, Global Labor Organization (GLO).

Bell, D. N. and Blanchflower, D. G. (2011). Young People and the Great Recession. IZA Discussion Papers 5674, Institute of Labor Economics (IZA).

Berghammer, C. and Adserà, A. (2022). Growing inequality during the Great Recession: Labour market institutions and the education gap in unemployment across Europe and in the United States. *Acta Sociologica*, 65(4):374–397.

Berson, C. and Botelho, V. (2023). Record labour participation: Workforce gets older, better educated and more female. Ecb blog, European Central Bank.

Blanchard, O. and Katz, L. (1992). Regional evolutions. *Brookings Papers on Economic Activity*, 23(1):1–76.

Blanchard, O. J. and Summers, L. H. (1986). Hysteresis and the European unemployment problem. *NBER Macroeconomics Annual*, 1:15–90.

Bodnár, K. (2018). Labour supply and employment growth. *ECB Economic Bulletin Articles*, 1.

Borusyak, K., Hull, P., and Jaravel, X. (2022). Quasi-experimental shift-share research designs. *Review of Economic Studies*, 89(1):181–213.

Bound, J. and Holzer, H. (2000). Demand shifts, population adjustments, and labor market outcomes during the 1980s. *Journal of Labor Economics*, 18(1):20–54.

Bratsberg, B., Raaum, O., and Røed, K. (2016). Job Loss and Immigrant Labor Market Performance. RF Berlin - CReAM Discussion Paper Series 1602, Rockwool Foundation Berlin (RF Berlin) - Centre for Research and Analysis of Migration (CReAM).

Bräuer, R. and Kersting, F. (2023). Trade shocks, labour markets and migration in the first globalisation. *The Economic Journal*, 134(657):135–164.

Bredtmann, J., Otten, S., and Rulff, C. (2017). Husband's unemployment and wife's labor supply: The added worker effect across europe. *ILR Review*, 71(5):1201–1231.

Cajner, T., Coglianese, J., and Montes, J. (2025). The long-lived cyclicality of the labor force participation rate. *Review of Economics and Statistics*. Forthcoming.

Cammeraat, E., Jongen, E., and Koning, P. (2023). The added-worker effect in the Netherlands before and during the Great Recession. *Review of Economics of the Household*, 21(1):217–243.

Canagarajah, S. (2011). Are migrants more likely than nationals to be unemployed during economic crisis? *World Bank*.

Caporale, G. M., Gil-Alana, L. A., and Trejo, P. V. (2021). Unemployment Persistence in Europe: Evidence from the 27 EU Countries. CESifo Working Paper Series 9392, CESifo.

Carlsson, M. and Eriksson, S. (2019). Age discrimination in hiring decisions: Evidence from a field experiment in the labor market. *Labour Economics*, 59:173–183.

Cavalleri, M. C., Luu, N., and Causa, O. (2021). Migration, housing and regional disparities: A gravity model of inter-regional migration with an application to selected OECD countries. OECD Economics Department Working Papers 1691, OECD Publishing.

Cedefop (2017). Investing in skills pays off: the economic and social cost of low-skilled adults in the EU. *Luxembourg: Publications Office. Cedefop research paper No. 60.*

Cipollone, A., Patacchini, E., and Vallanti, G. (2014). Female labour market participation in Europe: novel evidence on trends and shaping factors. *IZA Journal of European Labor Studies*, 3(1):1–40.

Cockx, B. and Ghirelli, C. (2016). Scars of recessions in a rigid labor market. *Labour Economics*, 41:162–176.

Coile, C. C., Levine, P. B., and McKnight, R. (2012). Recessions, Older Workers, and Longevity: How Long Are Recessions Good For Your Health? NBER Working Papers 18361, National Bureau of Economic Research, Inc.

Dao, M. C., Furceri, D., and Loungani, P. (2017). Regional labor market adjustment in the United States: Trend and cycle. *Review of Economics and Statistics*, 99(2):243–257.

Decressin, J. and Fatás, A. (1995). Regional labor market dynamics in Europe. *European Economic Review*, 39(9):1627–1655.

Dellas, H. and Koubi, V. (2003). Business cycles and schooling. *European Journal of Political Economy*, 19(4):843–859.

Engemann, K. M. and Wall, H. J. (2009). The effects of recessions across demographic groups. Working Papers 2009-052, Federal Reserve Bank of St. Louis.

Eurofound (2011). Shifts in the Job Structure in Europe During the Great Recession. *Dublin: European Foundation for the Improvement of Living and Working Conditions.*

Evans, R., Moore, A., and Rees, D. M. (2019). The Cyclical Behaviour of the Labour Force Participation Rate in Australia. *Australian Economic Review*, 52(1):94–106.

Fernández, R. (2013). Cultural Change as Learning: The Evolution of Female Labor Force Participation over a Century. *American Economic Review*, 103(1):472–500.

Furceri, D., Kilic Celik, S., Jalles, J. a. T., and Koloskova, K. (2021). Recessions and total factor productivity: Evidence from sectoral data. *Economic Modelling*, 94(C):130–138.

Galí, J. (2020). Insider-outsider labor markets, hysteresis, and monetary policy. NBER Working Papers 27385, National Bureau of Economic Research.

Goldsmith-Pinkham, P., Sorkin, I., and Swift, H. (2020). Bartik instruments: What, when, why, and how. *American Economic Review*, 110(8):2586–2624.

Gomes, D. B. P. (2024). Gender, Part-Time Work, and the Business Cycle. *International Monetary Fund*.

Hardoy, I. and Schøne, P. l. (2014). Displacement and household adaptation: insured by the spouse or the state? *Journal of Population Economics*, 27(3):683–703.

Hoynes, H. W., Miller, D. L., and Schaller, J. (2012). Who Suffers During Recessions? *NBER Working Papers 17951, National Bureau of Economic Research, Inc.*

Illing, H. and Koch, T. (2021). Who Suffers the Greatest Loss? Costs of Job Displacement for Migrants and Natives. IAB-Discussion Paper 202108, Institut für Arbeitsmarkt- und Berufsforschung (IAB).

Johnson, R. W. (2012). Older workers, retirement, and the great recession. In *The Russell Sage Foundation and The Stanford Center on Poverty and Inequality*.

Jordà, O. (2005). Estimation and inference of impulse responses by local projections. *American Economic Review*, 95(1):161–182.

Koehler, J., Laczko, F., Aghazarm, C., and Schad, J. (2011). Migration and the Economic Crisis: Implications for Policy in the European Union. *Research and Publications Division, IOM*.

Lee, G. H. and Parasnis, J. (2014). Discouraged workers in developed countries and added workers in developing countries? Unemployment rate and labour force participation. *Economic Modelling*, 41(C):90–98.

León, M. (2009). Gender equality and the european employment strategy: The work/family balance debate. *Social Policy and Society*, 8(2):197–209.

Li, D., Plagborg-Møller, M., and Wolf, C. K. (2024). Local projections vs. VARs: Lessons from thousands of DGPs. *Journal of Econometrics*, 244(2):105722.

Maestas, N., Mullen, K. J., and Powell, D. (2016). The Effect of Population Aging on Economic Growth, the Labor Force and Productivity. NBER Working Papers 22452, National Bureau of Economic Research, Inc.

Mattsson, P., Må nsson, J., and Greene, W. H. (2020). Tfp change and its components for swedish manufacturing firms during the 2008–2009 financial crisis. *Journal of Productivity Analysis*, 53(1):79–93.

Montiel Olea, J. L. and Plagborg-Møller, M. (2021). Local projection inference is simpler and more robust than you think. *Econometrica*, 89(4):1789–1823.

Notowidigdo, M. J. (2020). The incidence of local labor demand shocks. *Journal of Labor Economics*, 38(3):687–725.

OECD (2011). International Migration Outlook 2011. *OECD Publishing, Paris.*

OECD (2019). Working better with age, ageing and employment policies. *OECD Publishing, Paris.*

Oreopoulos, P., von Wachter, T., and Heisz, A. (2012). The short- and long-term career effects of graduating in a recession. *American Economic Journal: Applied Economics*, 4(1):1–29.

Périvier, H. (2014). Men and women during the economic crisis: employment trends in eight European countries. Sciences po publications, Sciences Po.

Pouliakas, K. and Branka, J. (2020). EU Jobs at Highest Risk of COVID-19 Social Distancing: Will the Pandemic Exacerbate Labour Market Divide? IZA Discussion Papers 13281, Institute of Labor Economics (IZA).

Prieto-Rodriguez, J. and Rodriguez-Gutierrez, C. (2003). Participation of married women in the european labor markets and the "added worker effect". *Journal of Socio-Economics*, 32(4):429–446.

Reyneri, E. and Fullin, G. (2011). Labour market penalties of new immigrants in new and old receiving west european countries. *International Migration*, 49:31–57.

Riedl, A. and Schoiswohl, F. (2015). Is there an added worker effect? – European labor supply during the crisis. *Focus on European Economic Integration*, (4):71–88.

Tuzemen, D. and Van Zandweghe, W. (2018). The Cyclical Behavior of Labor Force Participation. Research Working Paper RWP 18-8, Federal Reserve Bank of Kansas City.

Verdugo, G. and Allègre, G. (2020). Labour Force Participation and Job Polarization: Evidence from Europe during the Great Recession. IZA Discussion Papers 13425, Institute of Labor Economics (IZA).

Vlasblom, J. and Schippers, J. (2004). Increases in Female Labour Force Participation in Europe: Similarities and Differences. Working Papers 04-12, Utrecht School of Economics.

Wall, H. J. (2009). The "man-cession" of 2008–2009: It's big, but it's not great. *Federal Reserve Bank of St. Louis*.

Yagan, D. (2019). Employment hysteresis from the Great Recession. *Journal of Political Economy*, 127(5):2505–2558.

Zorlu, A. and Hartog, J. (2008). Employment Assimilation of Immigrants in the Netherlands. Tinbergen Institute Discussion Papers 08-057/3, Tinbergen Institute.

Appendix A: Additional Results

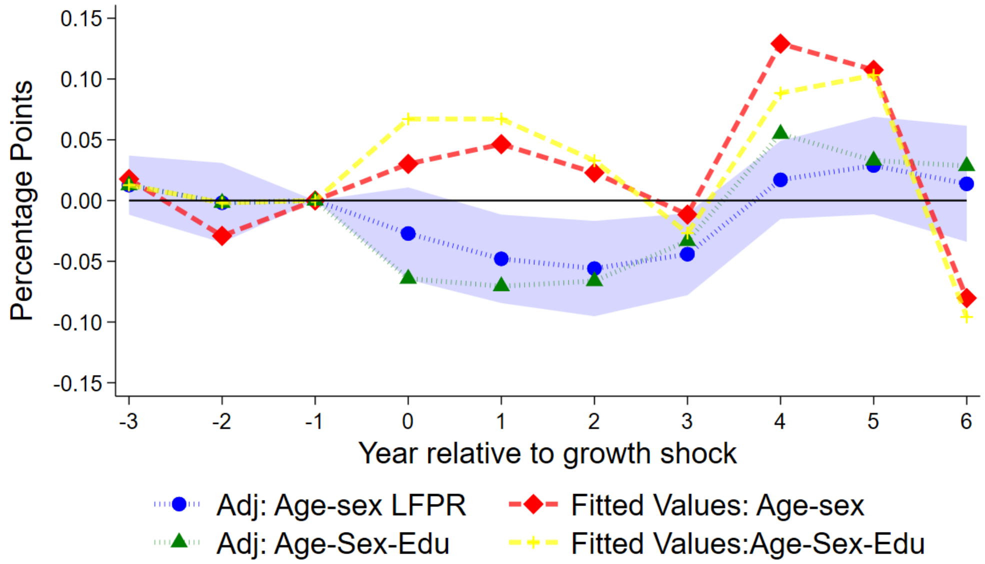

Figure 13: Estimated IRFs for adjusted age-sex LFPR and age-sex-education LFPR, along with their respective fitted values. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for k=0: 9.0.

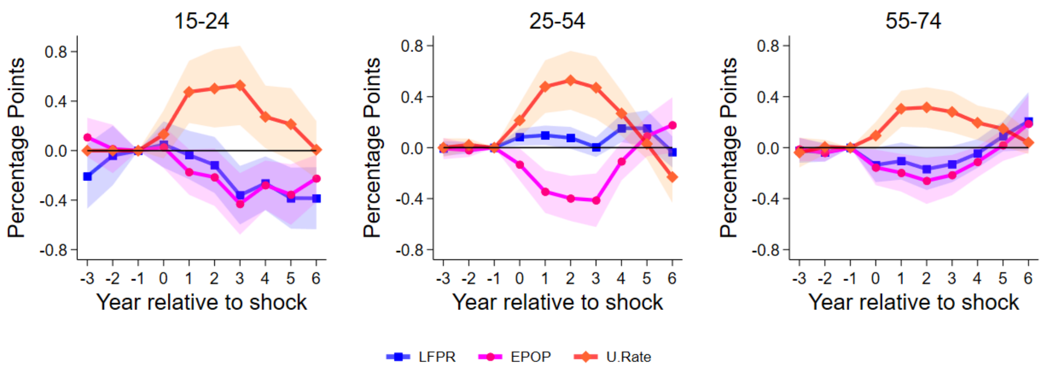

Figure 14: Estimated IRFs of the three **unadjusted** outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) for all age groups. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for $k = 0$: 16.0 for young, 13.7 for prime-aged and 17.4 for older workers.

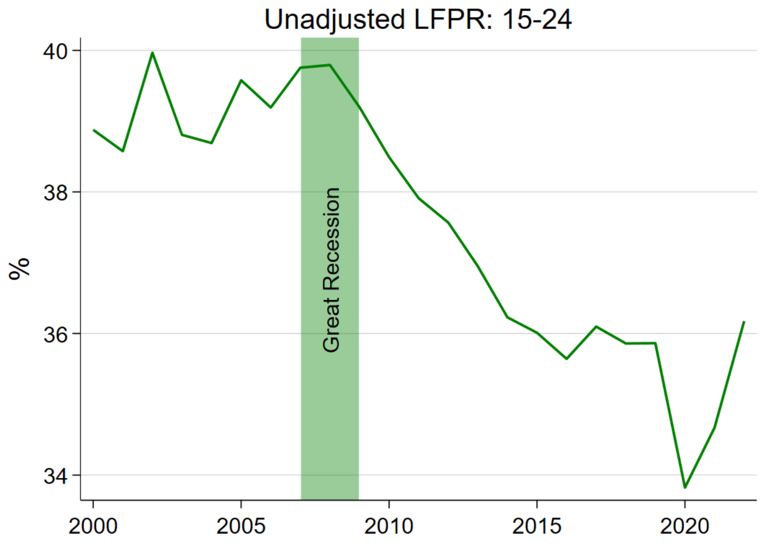

Figure 15: Unadjusted LFPR of the young (15–24). A marked decline begins around the Great Recession, with the pre-crisis level never fully recovered (descriptive evidence consistent with the hysteresis mechanism discussed in Section 5.3.2).

.

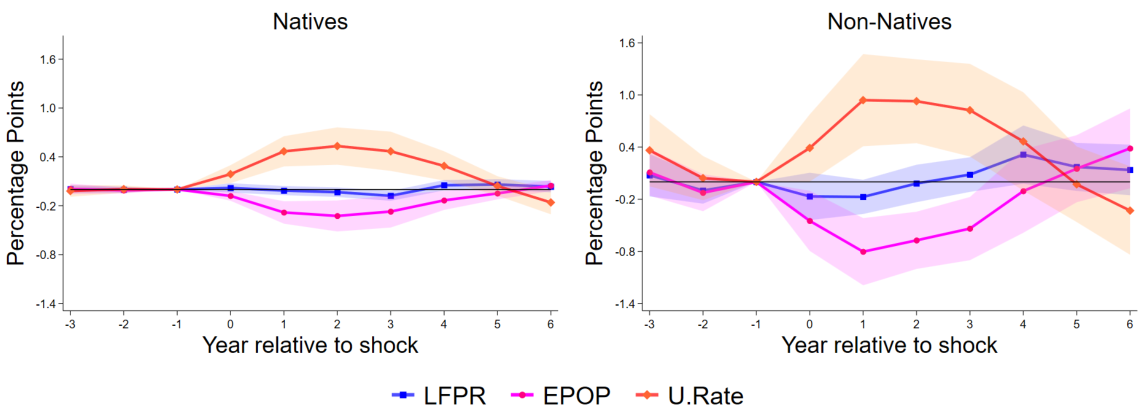

Figure 16: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) - by citizenship; **natives vs non-Natives (i.e. EU and non-EU foreigners)**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for $k = 0$: 16.8 for natives and 15.8 for non-natives.

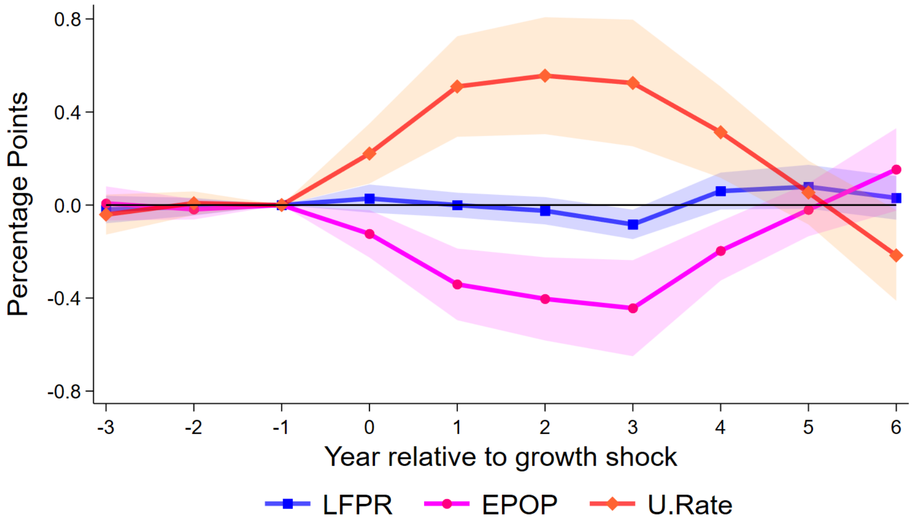

Figure 17: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) - **for the three-year lag structure of the industry shares ($\omega_{q,r,t-3}$)**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for $k = 0$: 14.5.

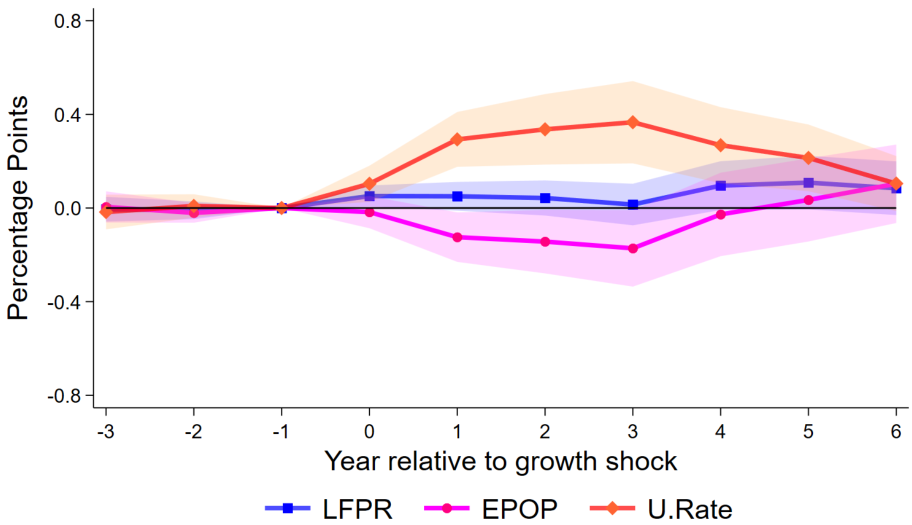

Figure 18: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) - **with no region-specific time trends included**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for $k = 0$: 38.12.

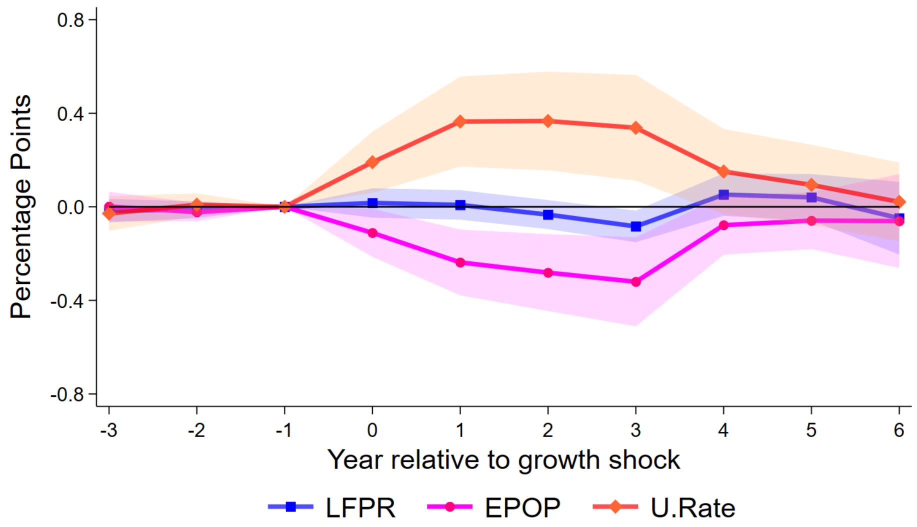

Figure 19: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) - **excluding year 2020**. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the regional level. F-statistic for $k = 0$: 14.5.

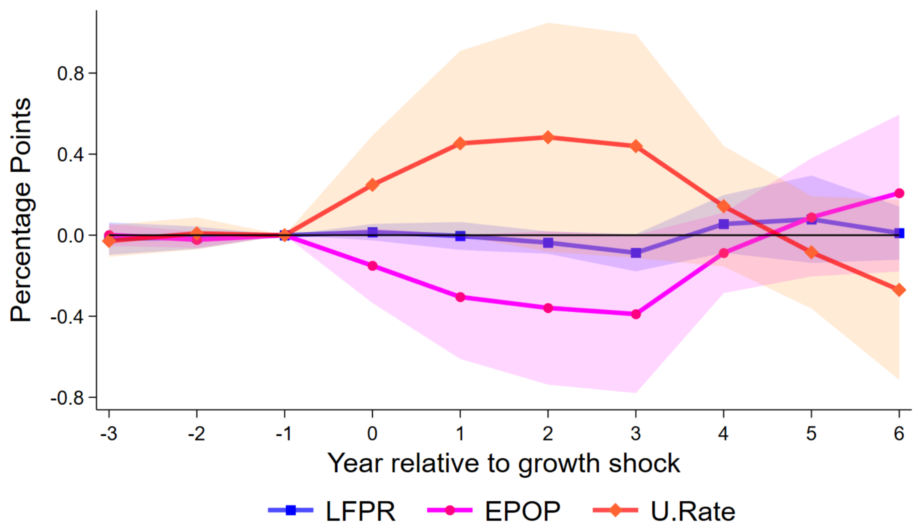

Figure 20: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange). Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing standard errors clustered at the **country level**. F-statistic for $k = 0$: 4.5.

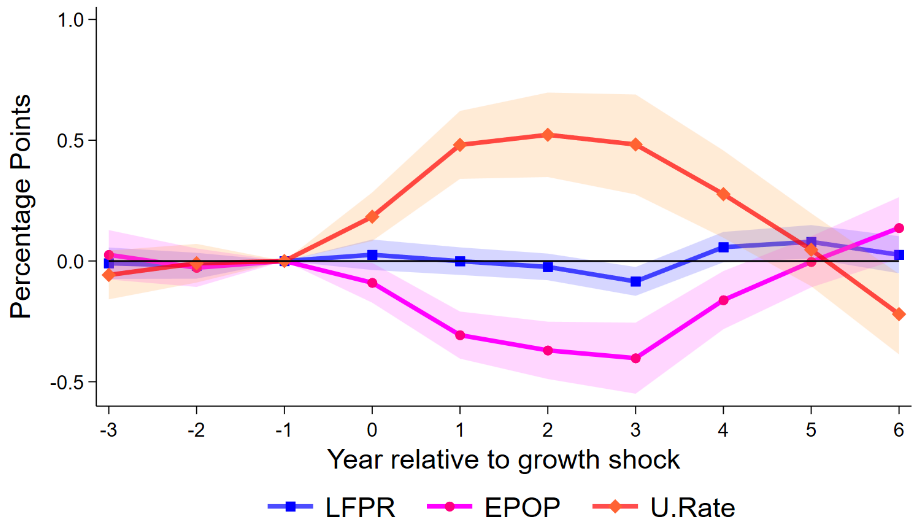

Figure 21: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange). Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, constructed using the shift-share inference procedure of Adão et al. (2019).

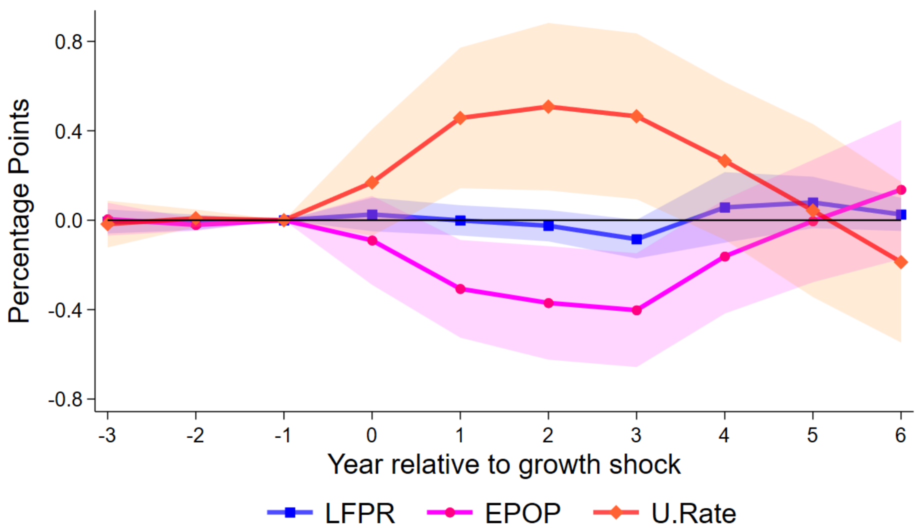

Figure 22: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange). Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing two-way clustering of standard errors by region and year. F-statistic for k=0: 4.

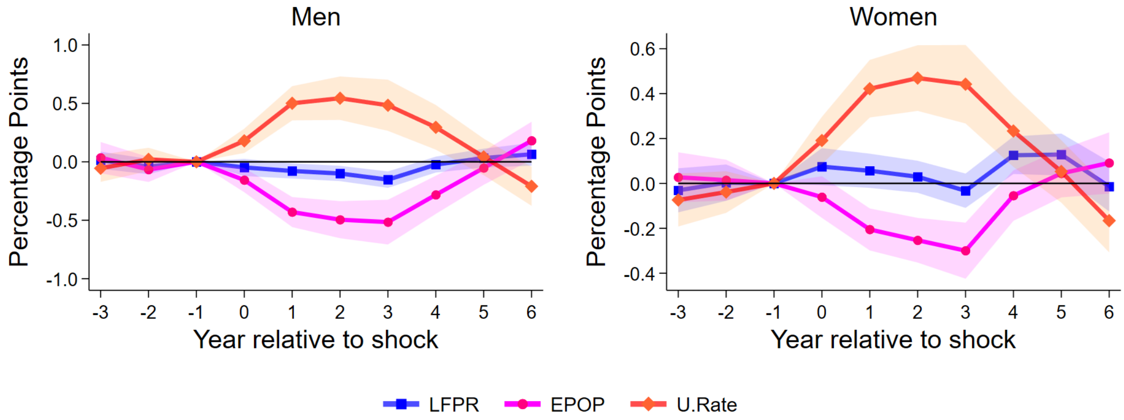

Figure 23: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) by gender. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, constructed using the shift-share inference procedure of Adão et al. (2019).

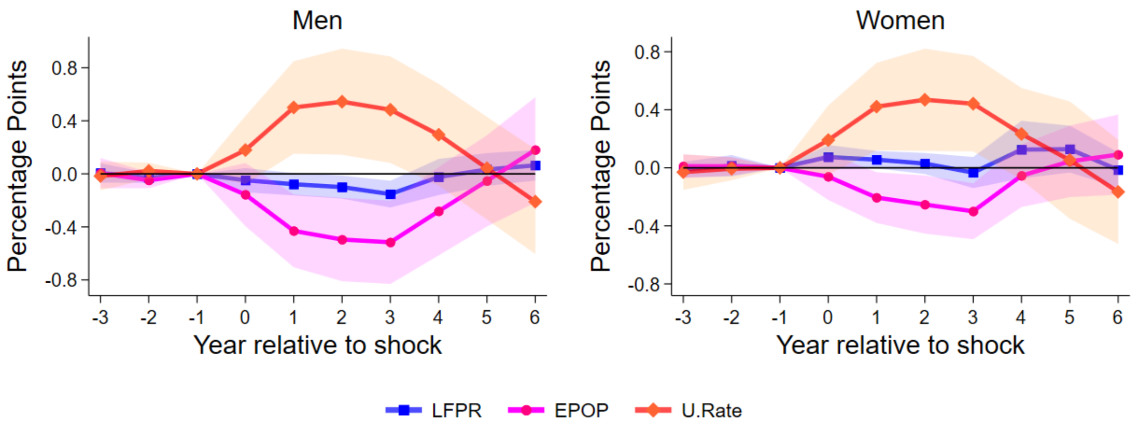

Figure 24: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) by gender. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing two-way clustering of standard errors by region and year. F-statistic for k=0: 4 for Men and 4 for Women.

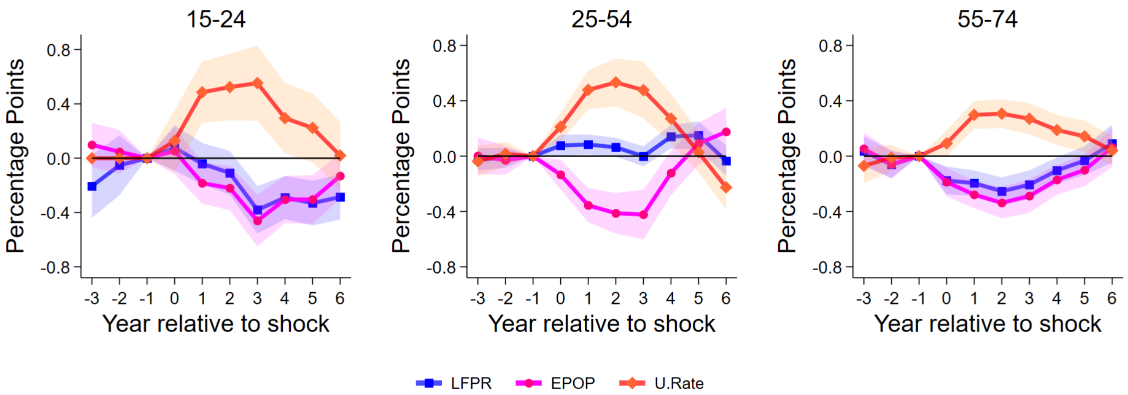

Figure 25: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) by age. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval,constructed using the shift-share inference procedure of Adão et al. (2019).

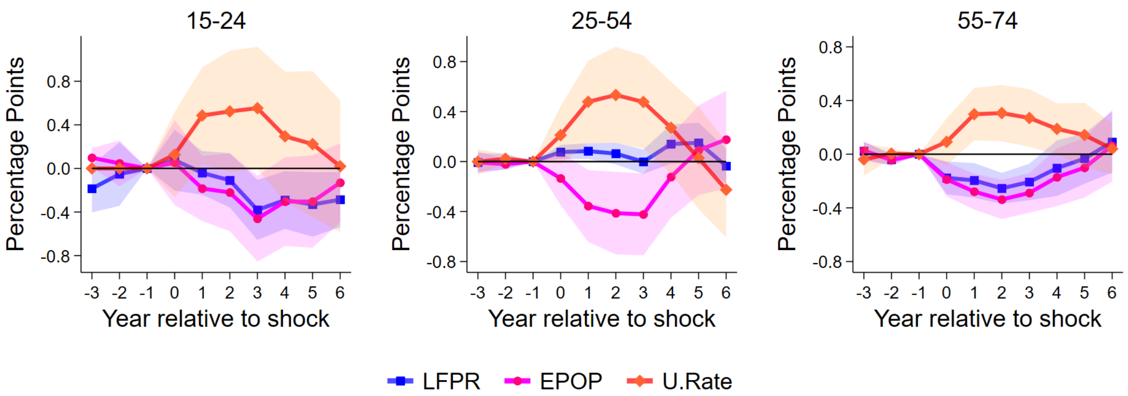

Figure 26: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) by age. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, representing two-way clustering of standard errors by region and year. F-statistic for k=0: 21.7 for 15-24, 4 for 25-54 and 4 for 55-74.

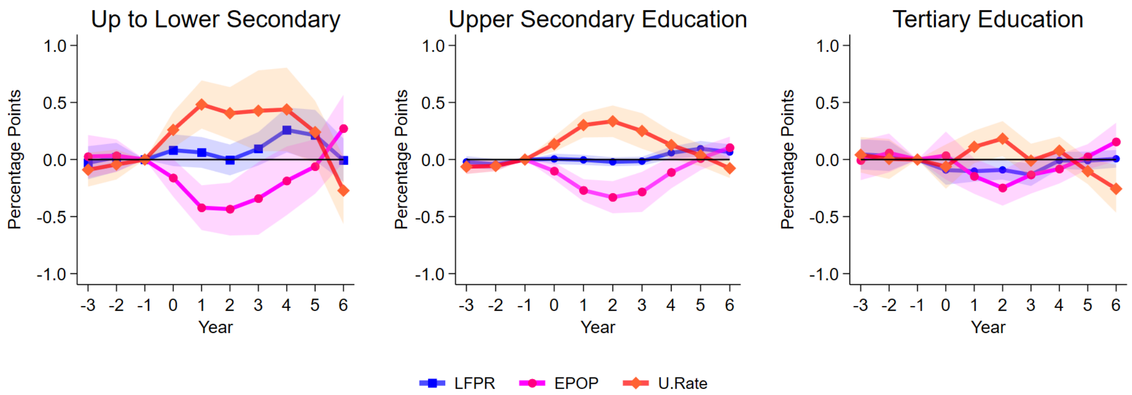

Figure 27: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) by education. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, constructed using the shift-share inference procedure of Adão et al. (2019).

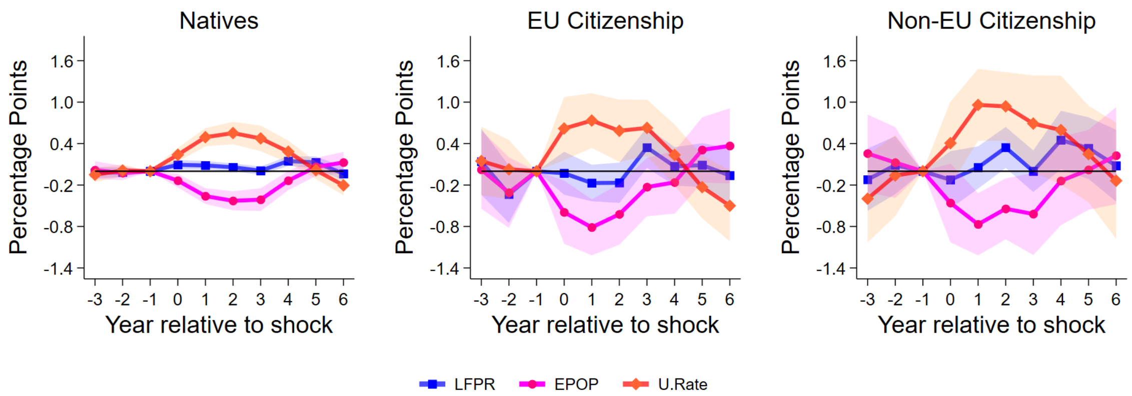

Figure 28: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) by citizenship. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval,constructed using the shift-share inference procedure of Adão et al. (2019).

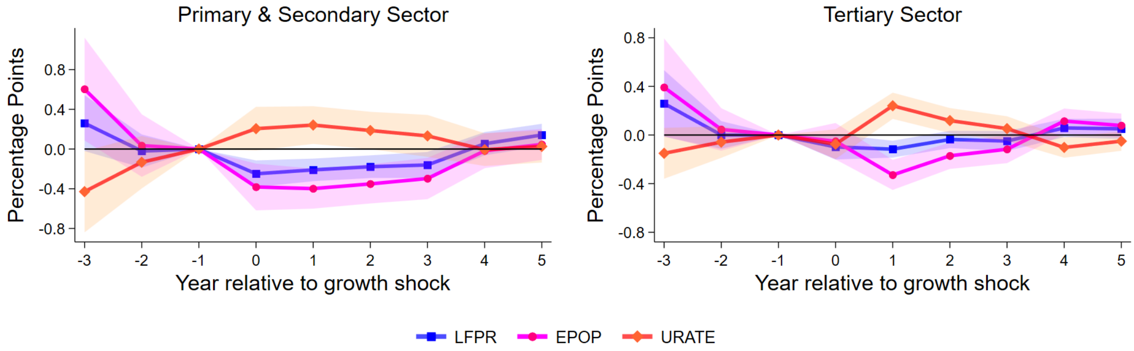

Figure 29: Estimated IRFs of the three age-sex adjusted outcome variables - LFPR (blue), Employment-to-Population Ratio (EPOP, pink) and unemployment rate (U.Rate, orange) by industry. Coefficients are normalised to show the effect of a temporary 1 percentage point decrease in output growth. The bands around each line indicate a 95% confidence interval, constructed using the shift-share inference procedure of Adão et al. (2019).

Industry	**Year**	α_{kt}	β_{kt}
Agriculture, Forestry & Fishing	2009	0.12	0.10
Industry (except construction)	2011	0.06	0.06
Construction	2006	0.06	-0.06
Professional, Scientific & Technical Activities	2011	0.05	0.04
Agriculture, forestry & fishing	2017	0.05	0.01
Construction	2010	0.05	0.03
Wholesale & Retail Trade	2019	0.05	0.03
Wholesale & Retail Trade	2012	0.03	-0.01
Industry (except construction)	2003	0.03	-0.02
Agriculture, Forestry & Fishing	2005	0.03	0.03
Total		0.52	0.42

Table 7: Rotemberg weights in GDP Shift-Share Instrument used in our main estimates. We show the 10 Rotemberg weights commencing from the third highest α_{kt}.

Acknowledgements

We thank Claudio Lissona, Paloma Lopez Garcia and Afonso Souto de Maura, the anonymous referee of the ECB Working Paper Series, as well as the participants of an ECB seminar for helpful comments.
The views expressed in this paper are those of the authors and do not necessarily reflect those of the European Central Bank (ECB).

Corinne Petrakis
Queen Mary University of London, London, United Kingdom; London School of Economics and Political Science, London, United Kingdom; email: a.c.petrakis@lse.ac.uk

David Sondermann
European Central Bank, Frankfurt am Main, Germany; email: david.sondermann@ecb.europa.eu

Postal address 60640 Frankfurt am Main, Germany
Telephone +49 69 1344 0
Website www.ecb.europa.eu

This paper can be downloaded without charge from www.ecb.europa.eu, from the Social Science Research Network electronic library or from RePEc: Research Papers in Economics. Information on all of the papers published in the ECB Working Paper Series can be found on the ECB's website.

PDF ISBN 978-92-899-7985-6 ISSN 1725-2806 doi:10.2866/5580297 QB-01-26-203-EN-N