Labour immigration and union strength

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Abstract

Is labour mobility in the European Union a threat to the strength of unions? We argue that the combination of cheap labour, workforce heterogeneity and low unionisation among labour immigrants is a potential challenge for unions. The challenge will be severe if immigration affects natives' unionisation. We use Norwegian administrative data in a natural experiment framework to examine this claim. The 2004 EU expansion led to a rapid increase in labour migration to the building and construction industry, but licensing demands protected some workers from the labour supply shock. We show negative labour market effects for workers exposed to labour immigration, but no effect on union membership. Our results question theories of unionisation, and are relevant for research on immigration, political behaviour and collective action.

KEYWORDS:
Immigration, Trade unions, Labour mobility, European Union

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Introduction

Migration of labour across countries can be a major force of change. Like international trade, it can create winners and losers, disrupt local and sectoral labour markets, and change domestic institutions. The impact of globalisation and Europeanisation on domestic institutions was a vibrant research area in the late 1990s (e.g. Cowles et al., 2001; Garrett, 1998; Rodrik, 1997) and has re-emerged with the intensification of international trade (Busemeyer, 2009) and the liberalization of labour mobility within the European Union (EU) (Afonso and Devitt, 2016). The consequences of a high supply of cheap labour (King and Rueda, 2008), earnings inequality (McCarty et al., 2006), and the dualization or segmentation of the labour market (Alt and Iversen, 2017; Emmenegger and Careja, 2012) are current political economy topics related to the consequences of labour migration. We study the consequences of increased labour mobility after the EU enlargement in 2004, a topic which has received massive attention both in the popular press and in academic work on Brexit and the rise of populism. (Colantone and Stanig, 2018; Finseraas et al., 2017)

The impact on trade unions holds an important position in the debates on the consequences of European labour migration. Although in decline, trade unions are still important agents in the political economy of many European countries. Norway, the case we study in this article, is a prime example. Unions and employer organizations are centralized at the national level, where coordinated wage bargaining determines wage growth intervals. Consultations between unions, employers and the government on relevant labour market issues are frequent, and it is politically difficult for the government to oppose agreements from previous tripartite consultations. This institutional model has been characterized as a stable, institutional equilibrium with support across the political spectrum. Indeed, some argue that this equilibrium is of key importance to the economic success of the Scandinavian economies (Barth et al., 2014).[1]

To what extent is labour mobility a threat to the institutional equilibrium in the labour market? While labour immigration can potentially have net positive, fiscal effects, the net benefit will be reduced/reversed if immigration has negative effects on well-functioning

institutions in the receiving country (Borjas, 2015). The rapid increase in labour immigration to Norway since 2004 has raised concerns regarding the future of the social model. These concerns are not only raised by the unions. Employer organizations and conservative politicians have also voiced concerns, which reflects the broad political support for the model.[2] We document that some of the concerns are real, as labour immigration has had important economic effects on parts of the Norwegian labour market. With this background, we use administrative register data to examine whether immigrant competition influences natives' propensity to unionise.

We highlight two potential effects of immigration on the motives to unionise. First, immigrants are less likely to unionise than natives, which might undermine the norm and social expectation of unionisation (Booth, 1985; Naylor and Cripps, 1993; Visser, 2002). Second, the willingness of workers to pay their membership fees should, on the margin, be influenced by whether they believe that the union can improve their situation in the labour market. Union density is one indicator of union strength, and is correlated with the existence of several types of pro-labour policies (e.g. Korpi, 2006) and the ability of unions to get their wage demands accepted (Wallerstein, 1989). Thus, falling union density in the industry due to immigration might weaken the instrumental incentive to unionise. We elaborate and critically assess these arguments in the next section, and also discuss the role of industrial relations traditions in Norway.

In order to identify the empirical effect of immigration on unionisation, we study the consequences of immigration in the Norwegian building and construction industry (BaC) after the 2004 EU enlargement. The enlargement led to a rapid increase in labour immigration from (in particular) Poland, Lithuania and Latvia; however, many workers are essentially protected from labour immigration due to various licensing demands. The shock from labour immigration combined with the licensing demands implies that data from the construction industry can be used to answer the more general question of what institutional responses to expect from workers who are exposed to increased competition from large and sudden international migration shocks. We do so by constructing 'treatment' and comparison groups of workers in the same industry, but who are exposed to

the supply shock differently. Using differences-in-differences designs, we conduct analyses that together identify both the short-term and long-term effects on exposed workers.

Our results show that the immigrant supply shock had negative effects on the earnings growth and the probability of employment for workers who were not protected by licensing demands. However, we find no evidence, neither short-term or long-term, of falling union membership among workers who remained in the labour market. Nor do we find that the supply shock changed the propensity to unionise among workers who entered the BaC industry. We conclude that while unions should be concerned about low organizing among immigrants, labour immigration is not a key reason for the decline of union membership among natives. We elaborate on these points in the conclusion and relate them to the broader literature on diversity and collective action.

Immigration and union density

Workers have social and instrumental motives for joining unions (Visser, 2002). In this section, we discuss these motives and, in turn, spell out why labour immigration might influence the motives to become a union member. We begin with the social motive since there are stronger arguments in favour of an effect on this motive. Next, we discuss how the particularities of the Norwegian case might influence the operation of these motives and the external validity of the Norwegian case. Finally, we briefly review the scarce empirical literature on this topic.

Immigration and motives to unionise

We document below that immigrants are less likely to organize. There are economic and cultural reasons for immigrants' reluctance to join unions, but irrespective of the reasons, lower unionisation rates among immigrants imply that native workers exposed to immigrant competition will have a higher share of non-unionized co-workers. A large literature, building on Akerlof (1980), emphasizes the importance of organized co-workers to explain the decision to join a trade union (e.g. Booth, 1985; Ibsen et al., 2017; Naylor and Cripps, 1993; Visser, 2002). When a large share of co-workers are union members,

the social motive for joining the union is strong and free-riding has a reputation effect. Importantly, shocks to union density will be persistent and even reinforced over time, and can thus change institutional equilibrium. The social motive for joining unions has strong empirical support, as workers typically highlight the importance of social reasons for why they joined a union (Visser, 2002, 406), and can explain low unionisation among newly arrived immigrants. The weakening of the social norm or custom of unionisation due to the inflow of unorganized workers might influence natives' propensity to unionise as well.

A related argument emphasizes the potential negative impact of workforce diversity on collective action (e.g. Korpi and Shalev, 1979; Stephens, 1979). This argument, which is particularly prevalent in the American literature on unions (Ferguson, 2016), can be traced back to Marx, who discussed the negative impact of Irish Catholic workers on the organization of the British working class (e.g. Afonso and Devitt, 2016, 4). As King and Rueda (2008) discuss, the prevalence of cheap labour can increase the saliency of ethnic identities in place of occupational identities, thereby threatening collective action at the workplace. Several mechanisms can explain why collective action becomes more difficult; for example, communication problems, preference diversity, distrust or prejudice are mechanisms that are emphasized in the research on the challenges of ethnic diversity for collective action (e.g. Alesina and La Ferrara, 2000, 2002; Alesina et al., 2001).

The decline in union density might also weaken instrumental incentives to unionise. On the margin, the willingness to pay the membership fee is likely to (also) be a function of the expected material benefits from being a union member. Strong labour unions have used their political influence to push for generous social insurance and labour market regulation, which is to the benefit of labour (e.g. Korpi, 2006). In this literature, falling union density is interpreted as a decline in union strength (Garrett and Lange, 1986; Korpi and Shalev, 1979). Since the ability to influence employers and policies is a function of organizational strength, declining density will weaken workers' incentive to unionise. Moreover, the power of unions in wage bargaining will depend on the share of the workforce they represent (Ahlquist, 2017; Wallerstein, 1989). Labour immigrants who do not organize weaken union strength and, thus, native workers' instrumental incentive

to unionise. Therefore, the social and instrumental incentives might reinforce each other, if, as often argued, workforce heterogeneity is as an obstacle to worker unity and union organization (King and Rueda, 2008; Korpi and Shalev, 1979).

While these arguments dominate the literature, there is a potential case to be made for a positive effect of immigration on the instrumental incentives to unionise. Immigration constitutes a labour supply shock for workers with similar skills as the immigrants. In textbook models of labour markets, an increase in supply will (in the short run) reduce the relative earnings of workers with comparable skills (Borjas, 2003), and can potentially increase the risk of unemployment (Dustmann et al., 2017). Such effects have been empirically identified in the Norwegian labour market (Bratsberg and Raaum, 2012; Bratsberg et al., 2014; Finseraas et al., 2017). For a given level of union strength, labour market competition might therefore improve the instrumental incentives to unionise, as a higher risk of income loss makes workers more willing to pay the membership fee to receive the insurance that unions provide (Blanchflower et al., 1990). A fall in absolute wages will work in the opposite direction if demand for union protection is a normal good. However, in line with standard economic demand and supply framework, we argue that immigration influences relative, not absolute, wages.

The argument of positive effects of competition on union membership is, however, controversial both theoretically and empirically. When competition for jobs is fierce, unions might be perceived as less effective, and their demands to employers will be less credible (Ashenfelter and Pencavel, 1969). Simply put, when unemployment is increasing, employers are dealt better hands. In line with this argument, union density tends to be pro-cyclical; that is, it falls when unemployment increases (Schnabel, 2013). Moreover, Schnabel and Wagner (2005, 16) used individual-level panel data from Germany and found no relationship between previous unemployment experiences and the propensity to be a union member.

The Norwegian case

The discussion above leads us to expect that, in general, the social motive to unionise will be weakened by the inflow of unorganized, immigrant workers. To what extent do the particularities of the Norwegian case modify this expectation? And how useful is the Norwegian case for inference to the general European case?

In Norway, national labour organizations are strong and have a good reputation for protecting workers' rights. Union density is comparatively high but below 50% in the BaC industry (see Online appendix). The labour market in Norway is strongly regulated, which, according to some observers, should make our case less susceptible to the disruptive effects of labour mobility (Afonso and Devitt, 2016). Occupational shocks might be less consequential for unions than in countries where unions are weaker and do not have this track record (Schulze-Cleven and Weishaupt, 2015). Norway does not have a Ghent system of social insurance, which implies that the selective incentives to unionise are not strong, and unions in the Norwegian BaC industry are so-called 'open shop' unions, which means that wage agreements cover all employees in firms covered by a tariff agreement (see the Online appendix for a brief description of the wage-negotiations regime). Since there is no discrimination between union members and non-members on this account, the open shop model weakens the importance of instrumental incentives for being a union member.

Given this context, the Norwegian case might represent a lower bound effect of immigrant competition on union membership. From the qualitative case selection literature (Gerring, 2007, 2008), one might consider the Norwegian case a 'least likely case' for immigration to influence unionisation: if we find effects in this case, the mechanisms are likely to also operate elsewhere. Thus, a critical test like ours is potentially very useful, but the flip side is that null findings might not travel to other contexts.

However, the Norwegian case also has some characteristics that moderate the least likely case impression. Most importantly, the social motive might be particularly sensitive to immigration in a country that has historically been very homogeneous along ethnic, religious and linguistic lines. Unions have limited experience in handling heterogeneity,

and social communication between workers might be particularly difficult. Since union membership is close to 50%, the share of workers who are on the margin of joining a union might be high. Moreover, there are selective incentives to unionise, as union membership implies access to cheaper private insurance offers, and perhaps most important, access to legal advice and union resources when facing the sack or lay-offs. Finally, the labour market is not strongly regulated along all dimensions; for instance, employment protection is at the OECD average (OECD, 2013).

We therefore believe that while the Norwegian case might not be the most representative case to study, it Mahoney and Goertz (2004) would not be characterized as an irrelevant case. We argue the results will have external validity for European countries that have a combination of organized, regulated labour markets and high levels of labour immigration since the EU expansion.

Previous empirical studies

The empirical literature on the effect of immigration on union density is relatively small. The main reason for this is the scarcity of good individual-level data on union membership. Studies from the US tend to find negative effects of immigration and ethnic diversity on union density (e.g. Ferguson 2016, but see Burgoon et al. 2010), while the cross-national literature produces divergent correlations (compare e.g. Brady 2007 and Lee 2005). However, most of this literature lacks research designs to disentangle the effect of immigration from correlated factors. Antón et al. (2016) is the article most similar to ours in motivation. They studied the effect of immigration on unionisation using Austrian data. An instrumental variables approach was applied to a prior, geographical distribution of immigrants to determine immigrant share, and substantive negative effects of immigration on union density were estimated. The negative effects were not driven by natives leaving unions, but by changes in the composition of the workforce in firms that increased their immigrant share. As described in the next section, we provide a more comprehensive analysis by conducting a set of analyses using individual-level rather than aggregated data.

1 The EU expansion, licensing demands and labour immigration

Immigrant inflow to Norway increased substantially over the last 20 years. In the mid-1990s, the total gross inflow of immigrants was about 15,000 a year, while in 2012, inflows reached about 65,000. All types of immigration increased over this period, but the most important increase is labour immigration after 2004. Prior to 2004, labour immigration to Norway was fairly limited and quite stable from year to year. From 2004, the European Economic Area (EEA) was expanded with 10 new member countries, including Poland. The expansion led to a rapid increase in labour immigration from a couple of thousand in 2004 to about 25,000 in 2012. About one third of the labour immigrants that arrived after 2003 entered the BaC industry (own calculation), which thus experienced a positive labour supply shift. While a large share of labour immigrants from the 'old EU countries' tend to return home after some years, about 70% of immigrants from the new member states settle permanently in Norway (Bratsberg et al. 2017: 22).

We exploit licensing and certification demands to get exogenous variations in the immigration shock within the BaC industry. Occupational licensing occurs when the law (or insurance companies) requires that all workers in an occupation pursue specialized vocational education to execute the tasks that fall into their profession.[3] For example, insurance companies and public building inspectors demand that plumbing and electric work is performed by workers with proper credentials, nationally approved licenses are in place to operate heavy machinery, and particular certificates have been earned to handle dangerous materials or to install lighting and light-signalling systems for roads, railways, airfields and harbour facilities. Such credentials typically follow from the completion of the relevant vocational education, and similar types of education or licenses from abroad are typically not accepted. The implication is that workers who have completed a vocational education programme that entails licensing and certification are effectively protected from labour immigration.[4]

To illustrate the effect of licensing on labour immigration, we present some illustrative means for so-called 'business areas'. The BaC industry can be divided into 16 business areas or trades, defined by five-digit industry codes. Each trade tends to be dominated by

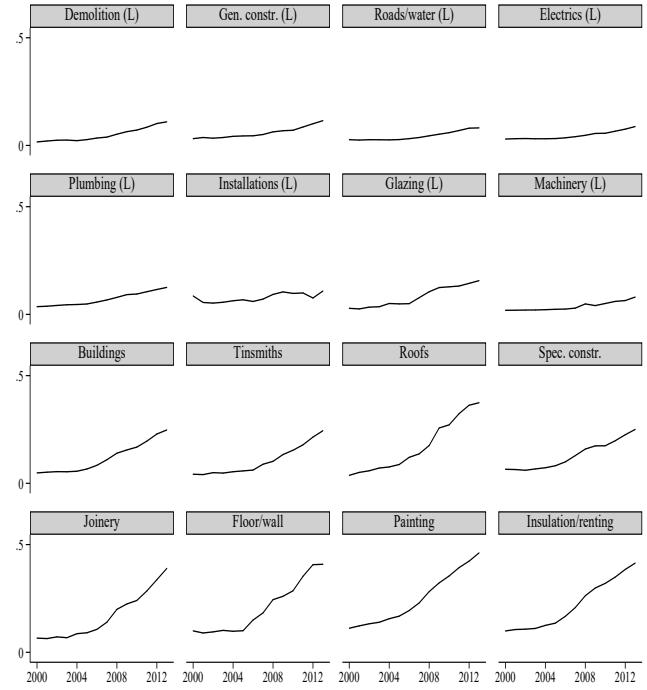

Figure 1: Immigrant share by trades
Note: Own calculations. All male workers in the BaC industry included. Workers' industry of employment is classified using employer-employee identifiers.

workers with one particular type of vocational education. Figure 1 graphs the development of the share of immigrants in each of the 16 trades from 2000 to 2013.[5] Trades that are dominated by protected workers are indicated by L in the figure. We see that the increase in immigrant share from 2004 is large in most of the trades without protection, compared to the trades with protection. In several of the non-licensed trades, the share of immigrants approaches 50% at the end of the period.

Labour immigrants are significantly less likely to unionise compared to natives. The Online appendix shows that the unadjusted immigrant-native gap across 2003–2013 is almost 14 percentage points. Moreover, the unionisation rate of immigrants is slow to converge to that of natives (Cools et al. 2018). Figure 2 visualizes the relationship between immigration and union density within trades over time. Each dot consists of 5% of the observations and is displayed in the figure according to the average union density and immigrant share within that bin, controlling for trade area and year fixed effects.

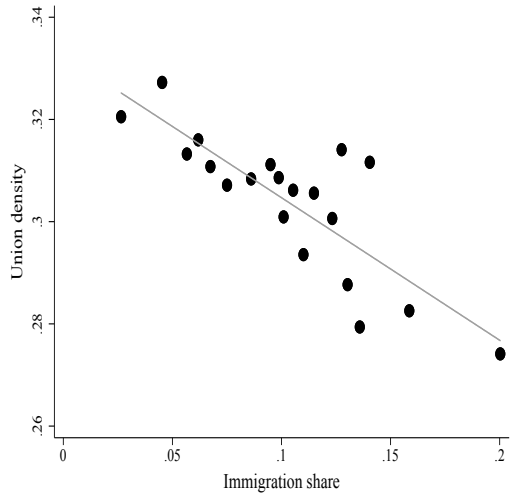

Figure 2: Immigrant share and union density
Note: The figure displays the relationship between immigrant share and union density, controlling for year and trade FE. The bins represent the mean union density for 20 equal-sized bins of immigrant share (see Stepner, 2013). The line is the regression line based on the underlying data.

The regression line is the relationship between immigrant share and union density based on the underlying data (not the bins). The figure shows a negative correlation between immigrant share and union density at the trade level.

Data, samples, and empirical strategy

The inflow of unorganized, immigrant labour constitutes the background for our study of how natives respond to immigration. We study the impact on natives' labour market outcomes and their propensity to unionise. As indicated above, the crux of our approach is to leverage the EU expansion and the licensing demands to compare construction workers who experienced the labour supply shock with a comparison group of construction workers who did not.

We use high-quality individual register data with a panel dimension, collected and organized by Statistics Norway. We conduct two types of analyses, using two different samples of individuals. In the first analysis, we follow workers in the BaC industry over time within their job spell. The sample in this analysis is all native males, 21–55 years of age, who are skilled workers and were employed in the BaC industry in the period 2000–

2013. We study males since the BaC industry is heavily male-dominated. By skilled workers, we refer to individuals who have fulfilled a vocational degree as their highest level of education. The analysis performed on this sample is explained below. In the second analysis, we follow skilled workers who were employed in the BaC industry in 2003, e.g. the year prior to the EU expansion. The second sample is all male native skilled workers, who were 21–55 years of age in 2003 and employed in the BaC industry that year. These individuals are followed on a year-by-year basis until 2013, irrespective of their employment status after 2003. The analysis performed on this sample is explained below.

We gather information on union membership from the administrative registers, which includes information on the annual union fees paid by the individual workers. This information is collected because a part of the union fee is tax deductable.[6] A worker is defined as a union member if the fee is positive in the respective year. The labour market outcomes are employment and log hourly wage. Employment is a binary variable taking the value of 1 if the individual is registered as a wage earner in the administrative employment register in the respective year. Log hourly wage is constructed from information on total wage payments in a given job, together with information on working time per week and number of days employed. When we follow *workers within their job spell*, employment and log hourly wage refer to employment and earnings in the BaC industry, while when we follow *workers employed in BaC in 2003*, employment and earnings can be from any industry in the economy. Descriptive statistics are reported in the Online appendix.

Exposed workers

For our empirical strategy to work, it is vital to identify (otherwise) comparable workers who are differently exposed to the immigration shock in their labour markets. For this purpose, we leverage the information about individual workers' education and occupational affiliation in the register data, combined with detailed information on licensing demands in different occupations, to identify exposed and protected workers. The key idea is to identify workers with vocational education who are in demand in the protected

occupations. In particular, we argue that native workers are protected from immigrant competition if they, through their type of vocational education, have access to the licensed occupations in the labour market. That is, for a worker to be protected by the regulations, the crucial factor is the extent to which the licensed part of the labour market is an available option for the worker.

In the Norwegian upper-secondary educational system, vocational programs consists of three or four years of education. Education from other countries may not fulfil the Norwegian requirements, or the foreign workers and/or their employers must go through a lengthy bureaucratic process to prove that they do. Thus, the legal requirements give native Norwegians a competitive advantage with regard to employment in licensed occupations. Thus, the legal requirements give native Norwegians a competitive advantage with regard to employment in licensed occupations, since immigrants cannot easily pursue this education. The Online appendix includes a detailed description of how the classification of vocational education programmes into the protected or exposed group is performed. From the classifications, we construct the variable E_i, which is equal to 1 if the worker i is exposed to immigrant competition and 0 if not. Throughout, our sample is restricted to workers who have completed vocational education in Norway. The exposed and protected groups are, by construction, similar with regard to years of education.

Follow workers within their job spell

Our first approach is to follow exposed and protected workers employed in the BaC industry over time. We estimate differences-in-differences (DD) models with the following structure:

$$y_{ijt} = \beta E_i * POST_t + \alpha_{ij} + \gamma_{ct} + \mu X_{ijt}^{'} + \epsilon_{ijt} \tag{1}$$

where y_{ijt} is the outcome for the worker i in job j in year t. α_{ij} refers to fixed effects for job spells, while γ_{ct} are county-year fixed effects. $X_{ijt}^{'}$ refers to a vector of time-varying individual controls that are included in some specifications. E_i is the time-invariant indicator of whether the worker was exposed to the supply shock (the treatment group), while $POST_t$ is an indicator for the years after the EU expansion. β is the DD-estimate.

The inclusion of α_{ij} and γ_{ct} absorbs the constituent terms of the $E_i * POST_t$ variable. Standard errors are clustered on the vocational education group.

It is important to realize that β is identified from variations within workers' job spells. Moreover, it is evident that the job spell has to cover both the pre- and post-period in order to contribute to the DD estimate. Thus, the estimate reflects movement in y_{ijt} from the pre- to the post-period, and is not driven by changes in what type of workers enter the two groups. The estimate is, however, clearly affected by changes in exit patterns, for instance, if workers in the licensed group are less likely to change job or leave the labour market. Although changes in exit patterns are a type of variation we want to capture, we are concerned that longer spells for the licensed group are correlated with other characteristics of the workers, such as age, seniority and experience. If so, β will be biased if these variables are omitted. We will therefore examine how sensitive the estimate is to such controls. In addition, we include interactions between E_i and an indictor for the pre-treatment years 2000–2002 to examine how sensitive the estimate is to potential deviations from the parallel trends assumption.

The estimates in equation 1 are essentially estimates of the effect of the EU expansion. In an extension, we use a similar approach to also derive estimates of the relationship between immigration share and the outcomes. More specifically, we estimate the effect of immigrant share using the following Instrumental Variable (IV) set-up:

$$y_{ijt} = \beta \widehat{IMSHARE}_{bt} + \alpha_{ij} + \gamma_{ct} + \mu X_{it}^{'} + \epsilon_{it}. \tag{2}$$

$\widehat{IMSHARE}_{bt}$ is the predicted immigrant share in trade area b.[7] The predicted share is from a first stage with $PIMFLOW_{bt} = IMFLOW_t * \frac{ELIS_{b,t=2003}}{ELIS_{t=2003}}$ as the instrument, where $IMFLOW_t$ is the inflow of immigrants to the BaC industry, $ELIS_{b,t=2003}$ is the number of workers in licensed occupations in trade area b in 2003, and $ELIS_{t=2003}$ is the total number of workers in licensed occupations in 2003. That is, we construct a predicted immigrant inflow by distributing all incoming immigrants to the BaC industry as if the initial licensing share of each trade completely determines the allocation of the incoming immigrants.

Our instrument is a variant of the widely used shift share instrument (Bartik, 1991). The key identifying assumption is that the initial shares of licensed workers in the trades are exogenous, conditional on the covariates. This assumption is plausible to the extent that the licensing and certification demands are in place mainly due to safety concerns and not to protect the labour market situation of the workers in those trades (e.g. Bratsberg and Raasum 2012). To interpret β from equation 2 as the effect of immigrant share, we need a set of additional assumptions (Angrist et al., 1996); in particular, that the only impact of the licensing demands over this period is running through immigrant share. Since this assumption might be violated, the estimate should be interpreted with caution; yet, we include it to provide an estimate of the structural parameter (immigrant share) of interest.

Follow workers employed in BaC in 2003

The second approach is to follow all workers employed in the BaC industry in 2003 on a year-by-year basis until 2013. We follow these workers irrespective of their choices after 2003, the idea being that all decisions after 2003 might be endogenous to the immigration shock. Our regression models have the following form:

$$y_{it} = \gamma_t + \delta E_i + \sum_{t=2000}^{t=2013} \beta_t E_i T_t + \epsilon_{it} \tag{3}$$

where y_{it} is the outcome for worker i in year t, γ_t are year fixed effects, and E_i is the exposure indicator. The β_t coefficients capture, on a year-by-year basis, how workers who were in the licensing area in 2003 differ, on average, from the non-licensed workers pre- and post-2003.

The key assumption for this 'starting line'-approach to reflect the effect of the immigration shock is that the two groups would have had a similar year-by-year trajectory without the EU expansion. This is a strong assumption. One concern is that a different composition of the two groups with respect to background characteristics will violate this assumption, for instance, if there is an age difference between the two groups. We adjust for initial differences between the two groups by re-weighting the sample so that they are

similar, on average, across a large number of variables, including lagged outcomes and interactions between the variables (Hainmueller, 2012). In the Online appendix, we describe the approach in detail and show that the two groups are balanced when the weights are applied. With weights, we are more confident that divergent trajectories in the two groups after the EU expansion are not due to different initial characteristics.

While both approaches identify the effects of the supply shock on exposed workers, the job spell approach estimates short-run effects on workers within their job spells. These estimates are useful because they are closely related to what happens with unionisation at the workplace. However, as it is tied to job spells, the approach fails to account for total and longer-term effects of the shock. Therefore, the two approaches complement each other to provide a fuller picture of the consequences of the shock.

Empirical results

Follow workers within their job spell: DD

Table 1 presents the results from the job spell DD analysis in equation 1. Panel A displays the earnings results. The first column shows the baseline estimate, which shows that exposed workers experienced a negative wage development compared to the protected workers. According to our estimates, hourly earnings growth was, on average, about 1–2 per cen% lower for those exposed to immigrant competition. Reassuringly, the DD estimate does not move much when we allow the exposed and protected workers to have a different trend in the pre-period (and the pre-trend is not significant). The estimate is also robust to the inclusion of the vector of (statistically significant) controls (column 3). In the final column, we separate early and late observations in the post-expansion period. More specifically, the early period is defined as 2004 to 2008, while the late period is defined as 2009 to 2013. We find that the late period coefficient is much larger than the small coefficient for the early period; thus, exposed workers who remained in their jobs experienced the strongest decline in relative earnings. This result is also in line with the continuing inflow of labour immigrants over time.

Panel B shows the corresponding estimates for union membership. The results are easy

to summarize: across the models, we find no significant DD estimates, and the substantive size of the coefficients is small. In the final column, we find that exposed workers who remain in their jobs are somewhat more likely to unionise. This result clearly goes against the social custom hypothesis and theories on the importance of workforce heterogeneity, and it is more in line with theories emphasizing the importance of employment risk. However, the estimate is small, and we cannot reject the null hypothesis of no differences between the groups. We therefore conclude that the labour supply shock following the EU enlargement had no impact on the union membership of exposed workers who remained in their jobs.

Table 1: Fixed effects regressions.

	(1) Baseline model	(2) Add pre-trends	(3) Add controls	(4) Two periods
Panel A: Log of hourly earnings				
$E_i * POST_t$	-.013**	-.015**	-.013**	
	(.005)	(.007)	(.005)	
$E_i * EARLY_t$				-.002
				(.005)
$E_i * LATE_t$				-.037***
				(.008)
F added var.		0.1	149***	14***
F $E_i * EARLY_t$= $E_i * LATE_t$				26***
Observations	619,990	619,990	619,990	619,990
No. of job spells	119,740	119,740	119,740	119,740
Y: Mean (SD)	4.88 (.5)	4.88 (.5)	4.88 (.5)	4.88 (.5)
Panel B: Union member				
$E_i * POST_t$	.005	.008	.004	
	(.008)	(.005)	(.009)	
$E_i * EARLY_t$				.002
				(.004)
$E_i * LATE_t$				.010
				(.010)
F added var.		2	14***	2
F $E_i * EARLY_t$= $E_i * LATE_t$				2
Observations	619,990	619,990	619,990	619,990
No. of job spells	119,740	119,740	119,740	119,740
Y: Mean (SD)	.43 (.5)	.43 (.5)	.43 (.5)	.43 (.5)

Note: All models include controls for job spell and county-year fixed effects (FE). Robust standard errors adjusted for clustering on vocational education in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Follow workers within their job spell: IV

Table 2 presents the results from the IV analysis.[8] For comparison, we include the OLS estimates at the bottom of the table. They show a negative correlation between immigrant share and earnings, and a weak and insignificant correlation for union membership. We expect the earnings estimate to be biased towards zero, as it is plausible that immigrants are recruited to trades with a high demand for labour. To the extent that these trades

also have higher union density, there will also be a positive bias in the OLS estimate on unionisation.

The second stage estimates confirm our expectation with regard to the earnings equation. The immigrant share estimate decreases substantively to -.78. Our estimates suggest that a ten percent increase in immigrant share reduces earnings with about 1 percent. With regard to unionisation, however, we find that the second stage point estimate is larger, not smaller, compared to the OLS estimate. This result implies that immigrants are recruited for low union density trades. Again, this result is more in line with competition increasing the propensity to unionise, than it is with social custom theory. As above, however, the estimates for unionisation are not statistically significant.

Table 2: Instrumental variable regressions.

	Earnings	Unionisation
IV		
Second stage		
Im. share	-.78***	.04
	(.20)	(.12)
First stage		
Pred. im. inflow	.05***	.05***
	(.01)	(.01)
Reduced form		
Pred. im. inflow	-.04***	.002
	(.01)	(.007)
OLS	-.50*** (.04)	.01 (.10)
Kleibergen-Paap F	13	
Observations	509,469	
No. of job spells	103,725	
Y: Mean (SD)	4.94 (.5)	
Im.share: Mean (SD)	.10 (.1)	
Pr. imflow: Mean (SD)	.25 (.3)	

Note: All models include controls for job spell FE, county-year FE, age-squared, seniority, seniority-squared, experience and experience-squared. Robust standard errors adjusted for clustering on vocational education in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Follow workers employed in BaC in 2003

The results thus far suggest that immigrant competition did not change workers' propensity to unionise among those who remained in the same job. However, as discussed, the

above analyses only captures one part of the effect of immigration and misses the potential impact of immigration, which is due to exits from the job or the BaC industry. The labour market competition might push some workers into a worse job than their initial one, others might upgrade their skills to avoid the competition, while still other workers might leave the labour market altogether. Movements to more precarious work are likely to be associated with falling unionisation, while upgrading might imply increasing unionisation to the extent that they enter jobs where unionisation is more common. To better capture these dynamics, we turn to the analysis where we follow workers employed in the BaC industry in 2003 on a year-by-year basis.

The results from estimating equation 3 are displayed in Figure 3. The first to note is that the re-weighting of the sample, which ensures that the groups are balanced in 2003, has the consequence that the treatment and control groups are essentially balanced also in the pre-treatment years 2000–2002. Next, we see that for all outcomes, there is a negative development for the exposed group compared to the protected group: relative earnings, employment probability and probability of union membership declines. For earnings and employment, the differences are statistically significant and amount to about 19% (earnings) and 16% (employment) of the standard deviation in 2013. Thus, the relative differences are economically important.[9]

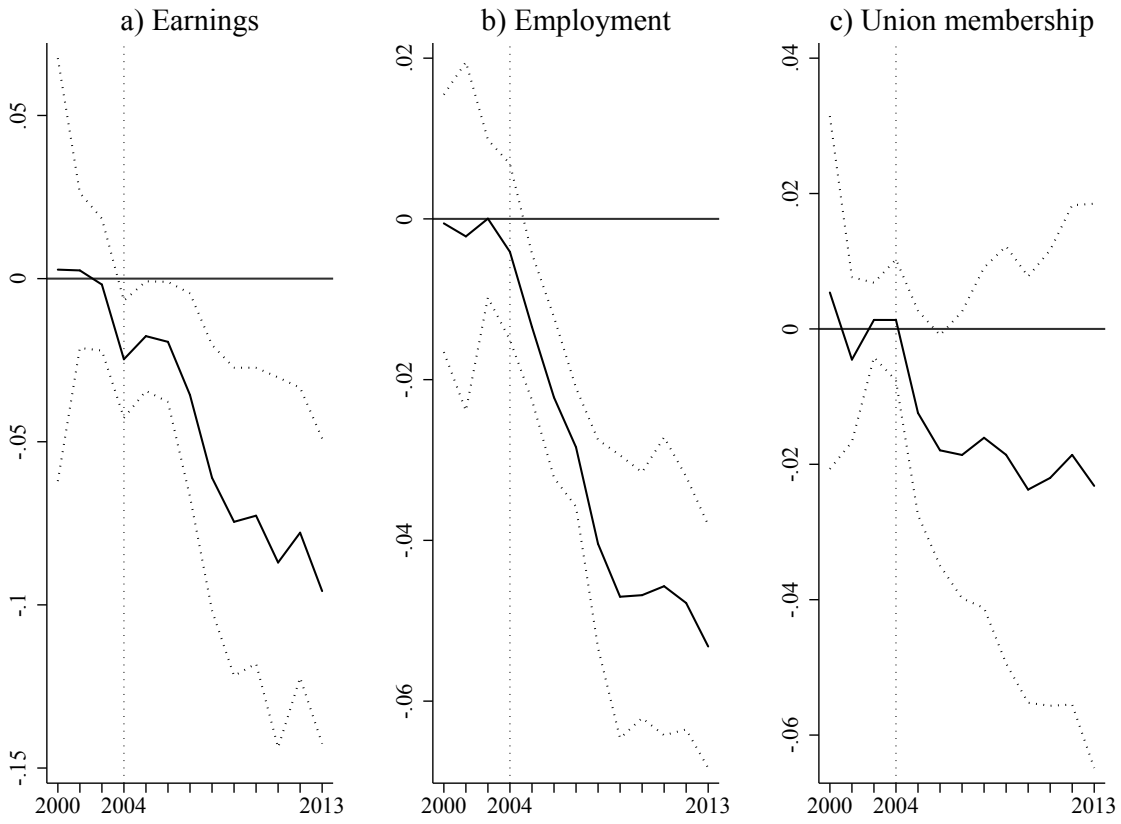

Figure 3: Year-by-year development in the gap between protected and unprotected workers.
Note: The figures plot the interaction terms between year and E_i (β_t) and the associated standard errors from the regression models in equation 3. All the year estimates are relative to 2003, for which the protected and unprotected groups are balanced using entropy balancing weights (Hainmueller 2012). The stippled vertical line indicates the year of the EU expansion.

For union membership, the difference in 2013 is smaller, about 5% of the standard deviation and with a p-value of .26. The increasing uncertainty in the unionisation trend over time that we see in the figure suggests that for many workers, joining a union is almost a one-time decision that is resistant to changing circumstances. Those who were union members at least once over the studied period were union members in about 70% of their observations. Thus, despite stability, there is meaningful variation over time. However, since most workers remain union members once they join, much of the variation over time comes from workers entering unions. If we restrict the analysis to workers who were not union members in 2003, we find a negative but not significant pattern in union

membership. The pattern is quite similar to the pattern for union membership in Figure 3. In any case, the results reinforce the conclusion of no support for the social custom hypothesis. In the Online appendix, we show that we get similar conclusions, although with more statistical precision, if we study cumulative outcomes over the years 2003–2013 instead of the year-to-year variations. Finally, we show that an analysis of unionisation among entrants further corroborates the finding that the supply shock had limited effects on natives' propensity to unionise.

Conclusion

The consequences of labour mobility have been a hot topic in several European countries that received a large increase in labour immigration after the 2004 EU enlargement (Colantone and Stanig 2018; Finseraas et al. 2017; Ruhs 2017). To what extent is the increase in labour mobility in Europe a threat to the organization of workers in the market sphere? We show that the increase in labour supply due to the EU enlargement had negative effects on the earnings and employment prospects of workers facing tougher labour market competition. However, we find no evidence that the increase in immigrant labour had any effects on natives' tendency to unionise. Our results are surprising in light of social custom theory: exposed workers experienced a rapid increase of unorganized immigrant co-workers into their labour market, which had significant economic impacts. The immigrants increase the share of unorganized co-workers, as well as the religious, cultural and linguistic diversity in these labour markets. A large literature argues that collective action, such as organizing workers, will be more difficult in such settings (e.g. Korpi and Shalev 1979; King and Rueda 2008). We find, however, no impact on the willingness to unionise.

While our study relies on appropriate data and a high level of internal validity, the external validity is harder to assess. One the one hand, the strong historical position of unions in wage-setting and policy-making at the national level might imply that the social norm of unionisation is strong despite negative shocks. If so, Norway is a 'least likely case' and the results might be a lower bound estimate of the effect of immigrant compe-

tition in a cross-national perspective. On the other hand, Norway has historically been ethnically and religiously homogeneous, which might make social norms more sensitive to immigration and unions less able to handle the immigration inflow.

We believe that our results are important for several literatures. First, our results speak to the literature on immigration, cheap labour, and the organization of coordinated labour markets (Alt and Iversen, 2017; Emmenegger and Careja, 2012; King and Rueda, 2008). The labour supply shock constitutes the introduction of cheap labour into a well-organized labour market and illustrates that it can have important economic consequences also in this context. While union membership among natives is unaffected, unions still face the challenge of organizing the newcomers. As Cools et al. (2018) showed, immigrants' unionisation slowly catches up with that of the natives with years since arrival. The slow catch-up process implies that unions might want to spend more resources on recruitment policies that target these groups. King and Rueda (2008) seem to favour union organization along ethnic and geographical lines as a response to non-organized immigrant labour. While we are skeptical of this solution, we believe that empirical evaluations of successful union strategies and policies to organize immigrants is a topic ripe for research, in particular, since it relates to the issue of immigrant integration into host societies more generally.

Second, the results are directly relevant for the social customs literature (Booth, 1985; Naylor and Cripps, 1993; Visser, 2002). While the existing literature on union membership tends to rely on cross-sectional regression analyses of samples of workers, we have population-wide panel data and rely on quasi-experimental variations in union organization. Our design is less susceptible to conflate the impact of organized co-workers with correlated characteristics of firms or industries, which is a serious concern in most of the existing research. Our results show that union members' willingness to pay their membership fees is quite resistant to changing circumstances in their industry of work, which is difficult to analyse with the type of data typically employed in the previous literature.

Finally, we consider our results as useful for the broader literature on ethnic diversity and political behaviour. In some respect, the willingness to pay the union membership fee

can be considered a behavioural measure of social solidarity, since the direct individual benefit of union membership is somewhat limited in our case. A large literature discusses the effect of ethnic diversity on various related outcomes (see Stichnoth and Van der Straeten, 2013), often with the hypothesis that blue collar workers will be particularly sensitive to shocks in diversity. Our results suggest that such effects do not extend to union membership.

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Notes

[1]See Acemoglu et al. (2012) for a more critical view of the Scandinavian model. There is also a large comparative political science literature on the political economy consequences of unions and wage bargaining (see e.g. Korpi, 2006).

[2]See e.g. the speech by Gerd Kristiansen (2014), then leader of the Norwegian Confederation of Trade Unions (LO), quotes from Svein Oppegaard (VG, 2011) of the Confederation of Norwegian Enterprise (NHO), and quotes from Michael Tetzschner (Dagbladet, 2011) of the Conservative Party.

[3]Here we rely on the Norwegian Occupational Regulations Database, NORD (Alecu and Drange, 2016; Bol and Drange, 2017). We are grateful to Ida Drange for sharing their data set.

[4]One may worry that some occupations received licensing protection after 2003, as a direct consequence of the increase in labour supply. Since we define workers as protected or not based on licensing regulations in 2003, additional licensing protection after 2003 would not represent a threat to our identification strategy, however, we should not expect to find effects from the EU expansion if many additional licenses were introduced. Reassuringly, Alecu and Drange (2016) show that none of the education programmes in our study changed licensing/cerification demands after 2003, but one occupational license was introduced in the BaC industry (scaffolders). The conclusions below remain if we exclude these workers.

[5]The Online appendix includes a similar graph for union density.

[6]The union fee is a flat rate of earnings and is about 1.65% (in our data) across trades. Due to increases in the tax deduction, the real price of union membership declines over the period, but the decline was the same across trades, sectors and industries. The fee is deductable also for labour immigrants.

[7]Ideally, we want to construct a measure of predicted immigrant share for each type of vocational education rather than for the 16 trade areas (see Figure 1). However, because there is a non-negligible share of immigrants with missing information on education, we opt for a trade area share instead. The Online appendix includes a discussion on measurement error in the immigrant labour supply.

[8]This analysis is limited to the years 2003–2013 because the necessary information on occupations is missing pre-2003.

[9]While we interpret the pattern as reflecting the labour supply shock, an alternative interpretation is that workers in protected sectors are more likely to stay in their jobs because they have made important investments in licenses and certificates. Since all workers in our sample have completed a vocational education programme, we do not think this explanation is likely. Nonetheless, if the alternative interpretation is correct, we should see that those in the treatment group are more likely to leave the BaC industry for work elsewhere. In the Online appendix we show the probability of remaining in the BaC industry, conditional on employment. Contrary to the investment hypothesis, we see that those in the treatment group are more likely to remain in the BaC industry.

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Online Appendix:
Finseraas, Røed, and Schøne (2019) 'Labour Immigration and Union Strength',
European Union Politics

Wage negotiations in the BaC industry

Trade unions in Norway are generally so-called "open shop" unions, implying that wage agreements cover all employees at the firm, conditional on whether the firm is covered by a tariff agreement, irrespective of union member status.

The main form of wage agreement in the BaC industry is a settlement between federations ('Forbundsvise oppgjør'). These are negotiations between the main employers' and employees' organizations in the BaC industry ('Byggnæringens landsforbund' and 'Fellesforbundet'). Negotiations and settlements take place every year, but the main settlement is bi-annual. Negotiations are only on the federation level in the main settlement years; otherwise, negotiations are on the central level. Only BaC firms that have a tariff agreement are covered by the agreement. According to survey information in 2012, approximately 70% of BaC industry firms were covered by a tariff agreement. Agreements between the federations are followed by local negotiations.

In 2005, The General Application Act was introduced for the BaC industry. The purpose of the Act was to ensure that foreign employees' terms of wages and employment are equivalent to those of Norwegian employees and to prevent social dumping. The Act was first introduced in the five counties of Oslo, Akershus, Østfold, Buskerud and Vestfold. Then, Hordaland followed in 2006, before the whole country was covered in 2007.

Union density by trades

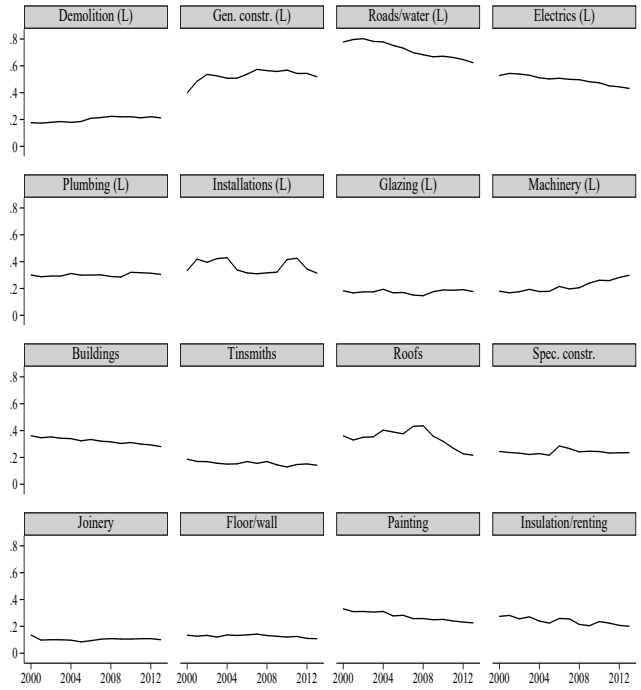

Figure A1: Union density by trades
Note: Own calculations. All male workers in the BaC industry included. Workers' industry of employment is classified using employer-employee identifiers.

Descriptive statistics

Table A1: Descriptive statistics.

	Obs.	Mean	St.dev.
Follow workers within their job spell			
E_i	619,990	.56	.49
Union member	619,990	.42	.49
Log hourly earnings	619,990	4.88	.49
Age	619,990	36.39	9.48
Seniority	619,990	5.96	5.77
Experience	619,990	17.38	9.48
immigrant share	619,990	.09	.07
Predicted inflow	619,990	.21	.28
Follow workers employed in constructionin 2003			
E_i	551,602	.55	.49
Union member	551,602	.46	.49
Log hourly earnings	551,602	4.87	.50
Employed	551,602	.91	21
Cumulative outcomes			
E_i	42,982	.56	.49
Union member	42,982	4.35	4.38
Log annual earnings	42,982	123	19.19
Employed	42,982	8.96	2.07
Δ imshare	42,982	-.00	7.84

The immigrant-native gap in union membership

Table A2: OLS regressions. Dependent variable is union membership.

	(1) Union member	(2) Union member	(3) Union member
Immigrant	-.14***	-.08***	-.07***
	(.04)	(.02)	(.02)
Trade FE	No	Yes	Yes
Year FE	No	No	Yes
N	1,689,346	1,689,346	1,689,346

Note: Robust standard errors adjusted for clustering on trade in parentheses.
*** p<0.01, ** p<0.05, * p<0.1.

The construction of E_i

To identify the effects of the EU enlargement and the subsequent immigration shock, we exploit that occupational licenses protect some of the skilled workers in the BaC industry from immigrant competition, while others are not protected by such regulations. The samples we study consist of skilled workers (see Section 4 for further details), defined as having a vocational education. Vocational education refers to having completed an educational programme in one of the occupation-oriented tracks in the upper-secondary level of the Norwegian education system. These tracks are standardized to four (sometimes three) years of schooling.[10] We restrict the sample to skilled workers because we can use their type of vocational education to determine if and to what degree individuals in this group have access to licensed occupations or not. Except for their access to licensed occupations, the skilled workers in the BaC industry are similar with regard to years of completed education and their exposure to business cycle fluctuations.

When an occupation is licensed, access to perform all (or vital) tasks within that occupation is regulated by laws that make unauthorized practice illegal or costly. The extent to which skilled workers have access to licensed occupations depends on their type of vocational education. We argue that native workers are protected from immigrant competition if they, through their type of education, have access to the licensed occupations in the labour market. That is, for a worker to be protected by the regulations, the crucial factor is to what extent the licensed part of the labour market is an available option, *not* that he (or she) is currently employed in a licensed occupation. Education from other countries may (often) either not fulfil the Norwegian requirements, or the foreign workers and/or their employers must go through a lengthy bureaucratic process to prove that they do.

To determine whether a skilled worker in the BaC industry is exposed ($E_i = 1$) to or protected ($E_i = 0$) from labour immigration, we proceed as follows:
First, we identify all (six-digit) vocational education codes held by skilled workers employed in the Norwegian BaC industry, 2003–2013. Then, we classify the 254 different vocational education programmes into 29 categories of relatively similar types of education; e.g., different kinds of electricians are grouped together and different types of plumbers are grouped together. This classification is listed and described with regard to the underlying codes in Table A3.
Second, we establish an indicator of the extent to which workers in each of the 29 groups of education have access to licensed occupations. For this purpose, we combine information about the occupational affiliation of all employed individuals in the register data with the Norwegian Occupational Regulations Database (NORD), in which occupations are defined as licensed or non-licensed (see Alecu and Drange 2016; Bol and Drange 2017). In these sources, occupations are classified according to the Norwegian standard of occupational classification (STYRK). The register data provides us with information about the STYRK code of all employed individuals, while the NORD database classifies occupations as licensed or non-licensed, using the same codes. Next, we use a sample of all employed individuals over the years of 2003–2013[11] and select workers between 18 and 55 years of age who hold one of the 254 NUS codes that we identified among the skilled workers in the BaC industry. Next, we calculate the mean share of employees who work in a licensed occupation within each of the 29 vocational groups. These mean shares, listed in Table 3A, are the indicators of the degree to which the skilled workers have access to the licensed

parts of the labour market.
Third, we define the individual workers in our analytical samples (described in Section 4) as exposed ($E_i = 1$) or protected ($E_i = 0$), based on the share of workers in the vocational group who work in a licensed occupation. We use the following rule: we define the individual worker as protected ($E_i = 0$) if the mean share of workers in the vocational educational group who work in a licensed occupation is greater than the overall mean share plus half of its standard deviation (see numbers in Table A3). If not, the individual worker is defined as exposed to competition ($E_i = 1$).

Table A3: Skilled BaC workers in licensed occupations, mean (M) and standard deviation (SD) of individuals hired in licensed occupations within educational category

	Educational categories	Education codes (NUS2000)	Mean	SD	Share of total emp. in BaC	E_i
1	Business and administration	440001–449999	4.4	20.7	5.61	1
2	Chemistry unspecified	452201–452299, 458403–458499	3.4	18.1	0.15	1
3	Information technology	454101–452299	6.0	23.7	0.39	1
4	Electrician	455101–455199	40.9	49.2	26.23	0
5	Mechanic	455201–455299	8.1	27.3	6.19	1
6	Precision mechanic	455301–455399	9.3	29.1	0.05	1
7	Asphalt, pavement	457102, 457199	31.7	46.5	0.81	1
8	Plumber	457113, 457121	60.5	48.9	8.18	0
9	Div. types of building education*	457101–457112, 457114–457120, 457122–457136	5.6	23.0	37.3	1
10	Construction vehicle/machinery operator	457901, 457999	55.6	49.7	7.09	0
11	Food processing	458101–458199	3.6	18.7	0.49	1
12	Textiles and footwear processing	458201–458299	4.4	20.5	0.09	1
13	Glazier/glass work	458300–458308	3.8	19.2	0.39	1
14	Instrument making and such	458308, 458312, 458314	12.4	33.0	0.01	1
15	Furniture and cabinet making, etc.	458309–458311, 458313, 458316–458329	6.2	24.2	0.60	1
16	Plastics mechanic	458315	6.1	24.0	0.08	1
17	Wood turning	458329	24.6	43.2	0.00	0
18	Mining and drilling	458401, 458402	30.6	46.1	0.08	1
19	Rock blasting, mining, stone work	458408, 458409	16.6	37.2	0.69	1
20	Tinsmith and other sheet metal work	459901–459999	10.0	30.0	0.84	1
21	Ambulance service	469901	75.8	42.8	0.04	0
22	Nursing assistance	461199–468999, 469902- 469999	37.8	48.5	1.05	1
23	Gardening	471101–473999	26.0	43.9	0.50	1
24	Forestry	474101–474999	12.3	32.9	0.27	1
25	Agriculture	479901–479999	12.7	33.3	0.76	1
26	Transport and navigation	481201, 481301, 481303, 481304, 481399	11.7	32.1	0.11	1
27	Professional driver	481401, 481499	57.5	49.4	1.60	1
28	Var. service and safety work	481901–489999	5.5	22.9	0.32	1
29	Unspecified vocational track	499999	8.3	27.6	0.07	1
Total		440001–499999	19.0	39.2	100	

Note: The means and SD are calculated from all employed individuals, 2003–2013, who hold one of the 254 NUS codes detected among the skilled workers in the BaC industry during the same period. See the above text in this Appendix for a detailed description. Shares of total employment in BaC are calculated from the second sample of skilled workers in the BaC industry, described in Section 4 in the main text. *Paperhanging, painting, concrete work, carpentry.

Entropy balancing of E_i

We conduct entropy balancing on union membership 2000–2003, log hourly earnings 2000–2003, log annual earnings 2000–2003, age, marriage, seniority, region of residence, and all possible interactions between these variables (except the lagged outcomes). We aimed for balance on two moments: means and variance. Table A4 shows the means in the treated ($E_i = 1$) and the control ($E_i = 0$) groups before entropy balancing, and the means in the control group after the construction of weights. Next, the table shows the standardized difference between the groups before and after balancing. As is evident, the sample is balanced on the means across all these variables when weights are applied. The re-weighting also achieves very good balance on the variances, the results of which are not included here.

Figure **??** shows the distribution of the entropy balancing weights (N = 43,603, mean = 1.12, median = 1, SD = .51, min = .00, max = 12.31). The number of large weights is small. We experimented with pruning these large weights, following Hainmueller's (2012) advice, and found that conclusions are very robust when doing so.

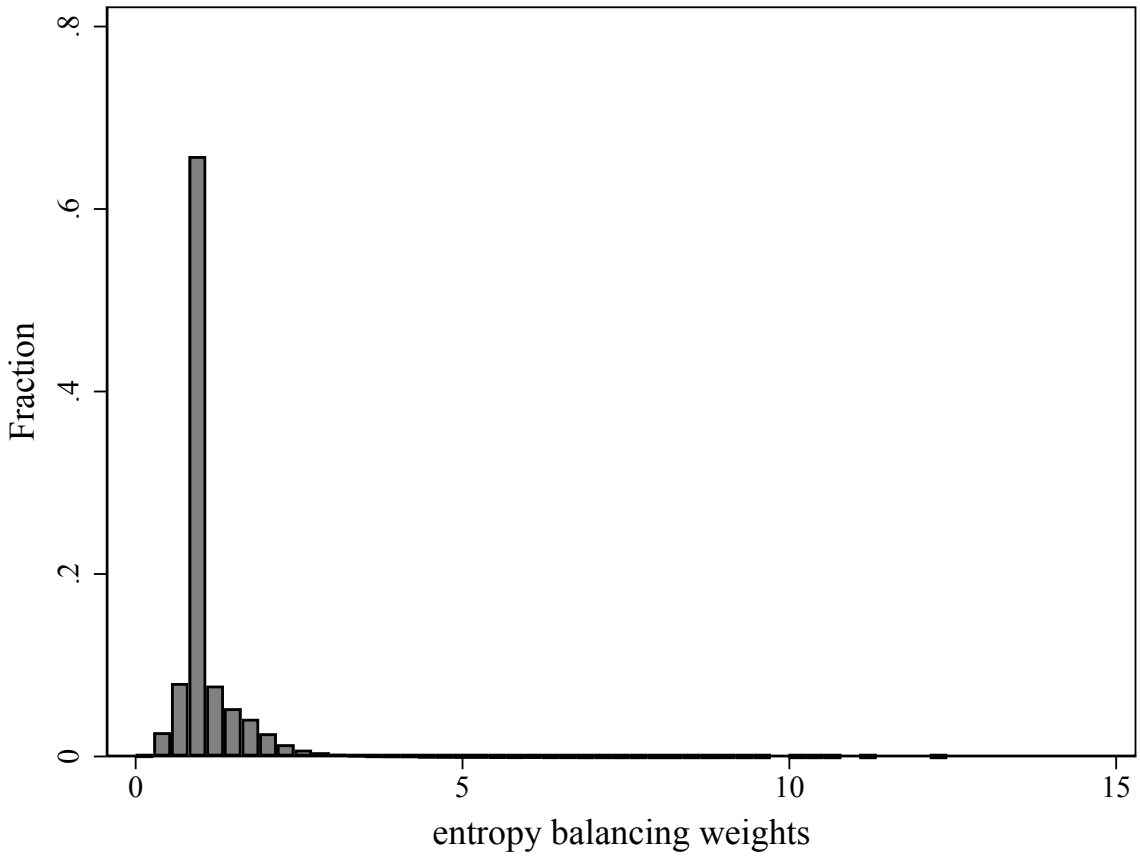

Figure A2: Histogram of entropy balancing weights.

Table A4: Balance pre and post entropy balancing.

	mean treated	mean control pre	mean control post	standardized difference pre	standardized difference post
Union member	0.40	0.54	0.40	-0.29	0.00
Union member 2000	0.41	0.53	0.41	-0.26	-0.00
Union member 2001	0.40	0.54	0.40	-0.28	-0.00
Union member 2002	0.40	0.54	0.40	-0.28	-0.00
Log hourly earnings	4.66	4.72	4.66	-0.11	0.00
Log hourly earnings 2000	3.98	3.93	3.98	0.04	0.00
Log hourly earnings 2001	4.12	4.37	4.12	-0.17	0.00
Log hourly earnings 2002	4.38	4.52	4.38	-0.12	0.00
Log annual earnings	12.64	12.69	12.64	-0.12	0.00
Log annual earnings 2000	12.08	12.33	12.08	-0.13	0.00
Log annual earnings 2001	12.26	12.48	12.26	-0.13	0.00
Log annual earnings 2002	12.46	12.61	12.46	-0.11	0.00
Age	35.79	36.03	35.79	-0.03	-0.00
Married	0.41	0.42	0.41	-0.01	-0.00
Seniority	4.87	5.42	4.87	-0.11	-0.00
Region 2	0.46	0.48	0.46	-0.05	-0.00
Region 2	1.89	2.58	1.89	-0.29	0.00
Region 3	167.55	170.37	167.54	-0.06	0.00
Region 4	22.97	25.80	22.97	-0.11	0.00
Region 5	1.94	1.98	1.94	-0.02	0.00
h. earningsXunion	0.50	0.45	0.50	0.04	0.00
h. earningsXage	1.31	1.23	1.31	0.04	0.00
h. earningsXseniority	2.13	2.27	2.13	-0.06	0.00
h. earningsXmarried	0.30	0.30	0.30	0.00	0.00
h. earningsXRegion 2	5.11	6.92	5.11	-0.29	0.00
h. earningsXRegion 3	453.08	457.65	453.06	-0.04	0.00
h. earningsXRegion 4	61.88	69.02	61.88	-0.11	0.00
h. earningsXRegion 55.22	5.29	5.22	-0.01	-0.00	
a. earningsXunion	1.37	1.20	1.36	0.04	0.00
a. earningsXage	3.53	3.30	3.53	0.04	0.00
a. earningsXseniority	5.76	6.08	5.76	-0.05	-0.00
a. earningsXmarried	0.82	0.80	0.82	0.01	0.00
a. earningsXRegion 2	15.00	20.26	15.00	-0.27	0.00
a. earningsXRegion 3	1.96	2.86	1.96	-0.22	-0.00
a. earningsXRegion 4	0.17	0.24	0.17	-0.18	0.00
a. earningsXRegion 5	0.05	0.06	0.05	-0.04	0.00
unionXage	0.10	0.12	0.10	-0.07	-0.00
unionXseniority	0.18	0.26	0.18	-0.20	0.00
unionXmarried	0.02	0.03	0.02	-0.10	-0.00
unionXRegion 2	190.17	210.72	190.19	-0.09	-0.00
unionXRegion 3	16.87	17.21	16.87	-0.02	-0.00
unionXRegion 4	3.81	3.49	3.81	0.03	0.00
unionXRegion 510.00	9.22	9.99	0.05	0.00	
ageXseniority	16.33	17.25	16.33	-0.05	-0.00
ageXmarried	2.23	2.27	2.23	-0.00	0.00
ageXregion 2	2.52	2.74	2.52	-0.05	-0.00
ageXregion 3	0.56	0.61	0.56	-0.02	0.00
ageXregion 4	1.48	1.49	1.48	-0.00	0.00
ageXregion 5	2.11	2.53	2.11	-0.10	-0.00
seniorityXmarried	0.33	0.30	0.33	0.01	0.00
seniorityXregion 2	0.04	0.04	0.04	0.01	0.00
seniorityXregion 3	0.13	0.12	0.13	0.03	0.00
seniorityXregion 4	0.18	0.19	0.18	-0.03	-0.00
seniorityXregion 5	0.03	0.03	0.03	-0.01	0.00
marriedXregion 2	0.11	0.09	0.11	0.04	0.00
marriedXregion 3	0.28	0.26	0.28	0.04	0.00
marriedXregion 4	0.46	0.48	0.46	-0.05	-0.00
marriedXregion 5	0.07	0.06	0.07	0.01	0.00

Measurement error in immigrant labour supply

Our measure of the immigrant labour supply may contain measurement errors for several reasons. First, a non-negligible portion of immigrant labour in recent years has been employed through temporary work agencies (TWAs). Many of these immigrants work in the BaC industry, although they are registered as workers in the TWA industry. In a recent report, Nergaard (2017) estimated that the share of TWA workers in the BaC industry is between 5 and 10%. Second, after the EU expansion in 2004, a relatively large share of immigrants in the BaC industry were hired by foreign contractors. That is, they worked in the Norwegian BaC industry, but they were employed by foreign firms (Dølvik and Eldring, 2008). As a consequence, they are not registered in the Norwegian employment registers.[12] Third, immigrants in the BaC industry, may work "off the books" to a different extent than native employees. In a survey of Polish BaC workers in 2010, 26% reported that they did not pay taxes (Eldring and Friberg, 2011).

All three possible sources of measurement error will most likely lead to an underestimation of the share of immigrants in the BaC industry. Most directly, it will affect the IV analyses. If registered and unregistered immigrants in the BaC industry are positively correlated, and if they both have the same effect on unionization, this will lead us to overstate the effect of immigration on unionization. Hanson (2006) discussed the distinction between legal and illegal immigration in the US setting, and argued that because the omitted variable in this case is immigrant-related, instead of classifying it as a form of measurement error, one could argue that the estimated effect is the total effect of immigration (both legal and illegal). Note that in the DD analyses, the measurement problems will be reduced to a large extent, since identification is not based on direct measurement of the labour supply of immigrants in the BaC industry.

Analysis of cumulative outcomes

In this section, we study the cumulative outcomes over the years 2003–2013 instead of the year-to-year variations. That is, we measure the cumulative yearly earnings, total years of employment, and total years of union membership. We regress these outcomes on i) the dummy for having an unprotected education in 2003 (Table A5 Panel A) and ii) on the change in immigration share in the trade of employment in 2003–2013 (Panel B). We centre the change in immigration share on its mean so that the constant in Panel B refers to the mean outcomes for those experiencing the average immigration supply shock. Like before, this approach produces statistically and economically significant effects on earnings and employment. Once again we find a negative and insignificant effect on union membership when we compare protected and unprotected workers. In Panel B, however, we find those who were employed in trades who experienced larger increases in immigrant share have fewer total years of union membership over this period. A one standard deviation increase in immigration amounts to a decline of about .8 year units, i.e. about 10 months of union membership. When we instrument the immigration shock using the licensing demands (Panel C), we find, however, no effect of immigration. Thus, the cumulative analysis provides the same conclusion as before with regard to the effect of immigrant competition.

Table A5: OLS regressions. Cumulative effects.

	(1) Earnings	(2) Emp	(3) Union
Panel A: Not protected			
Not protected	-3.300***	-0.348***	-0.262
	(0.639)	(0.044)	(0.569)
Constant	125.122***	9.124***	4.089***
	(0.469)	(0.030)	(0.535)
Y mean (st.d)	124 (19)	9 (2)	4 (4)
Observations	42,982	42,982	42,982
Panel B: Immigrant share			
Δ imshare	-0.246***	-0.025***	-0.100***
	(0.036)	(0.004)	(0.017)
Constant	123.361***	8.939***	3.914***
	(0.292)	(0.024)	(0.301)
Y mean (st.d)	124 (19)	9 (2)	4 (4)
X mean (st.d)	13(8)	13 (8)	13(8)
Observations	42,982	42,982	42,982
Panel C: 2SLS, second stage			
Δ imshare	-0.341***	-0.036***	-0.027
	(0.025)	(0.004)	(0.056)
Constant	123.319***	8.934***	3.946***
	(0.119)	(0.020)	(0.252)
Y mean (st.d)	124 (19)	9 (2)	4 (4)
X mean (st.d)	13(8)	13 (8)	13(8)
F first stage	34	34	34
Observations	42,982	42,982	42,982

Note: Robust standard errors adjusted for clustering on vocational education in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Propensity to remain in the BaC industry

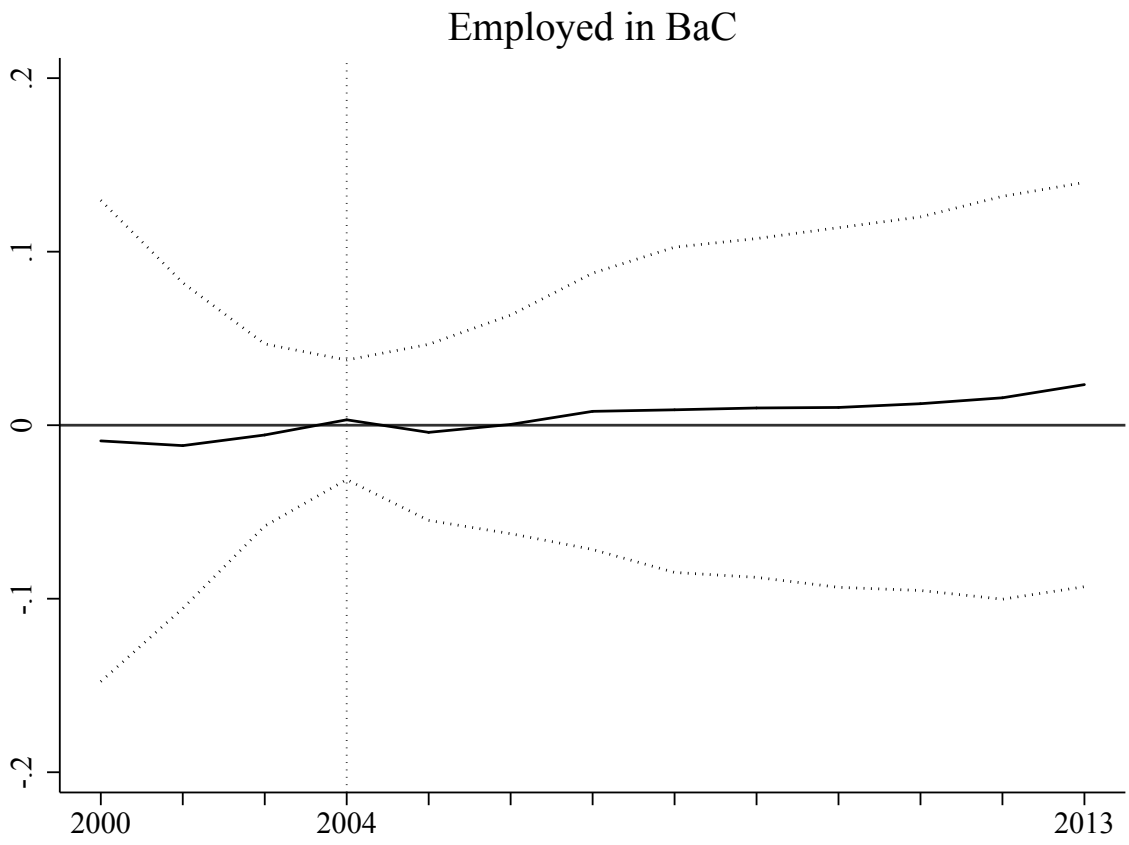

Figure A3: Year-by-year development in the gap between exposed and not exposed workers.

Note: The figures plot the interaction terms between year and E_i (β_t) and the associated standard errors from regression models of the form in equation 3. All the year estimates are relative to 2003, for which the exposed and not exposed groups are balanced using entropy balancing weights (Hainmueller 2012). The stippled vertical line indicates the first year in the post-treatment period.

Propensity to unionize among entrants

Ìn the main analyses we study the reaction of workers who were employed in the BaC industry prior to the shock. Here we address is whether the propensity to unionise differs among those who entered the BaC industry after 2003. Newly recruited workers' willingness to unionise might be more sensitive to the current context (Ibsen et al. 2017) than workers who were employed before the EU enlargement. To explore this issue, we study union membership rates in year $t+1$ for workers who entered the BaC industry in year t. We do so separately for the protected and exposed groups to examine whether the trends are different.

Figure A4 shows the trends in union membership for the two groups. We find a slight decline in union density over time. Throughout the period, new entrants who have undertaken protected education are more likely to unionise. More importantly, however, we see no change in the difference between the groups after 2004. This result further corroborates the finding that the supply shock had limited effects on natives' propensity to unionise.

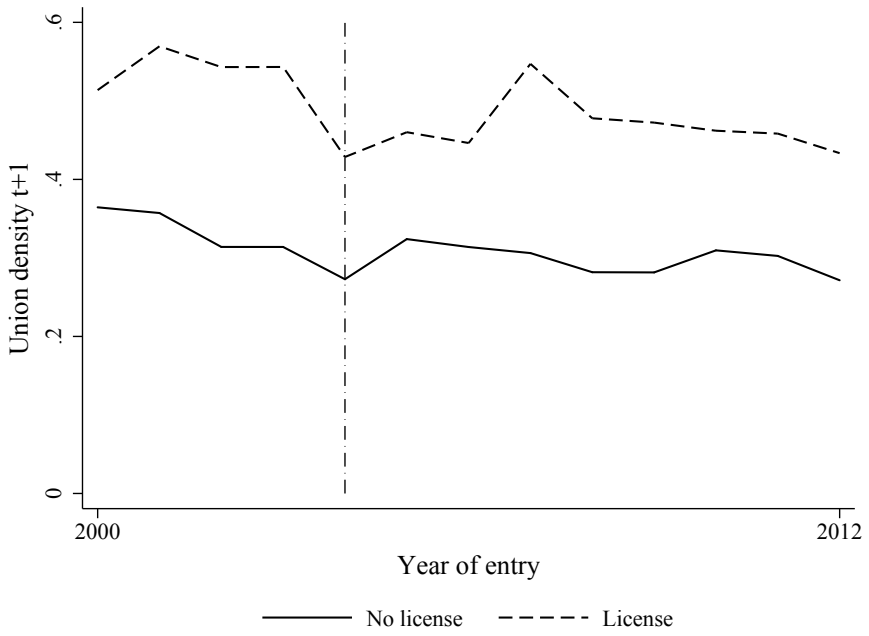

Figure A4: Trends in union density among entrants to the BaC industry.
Note: The full line shows the union density at year $t+1$ among exposed workers who entered the BaC sector in year t. The dotted line shows the corresponding trend for not exposed workers.

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