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## Workplace Segregation and the Labour Market Performance of Immigrants

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*Sébastien Willis*

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# Workplace Segregation and the Labour Market Performance of Immigrants

## Abstract

Immigrants are more likely to have conationals as colleagues, however the consequences of such workplace segregation is an open question. I study the effect of the conational share in an immigrant's first job on subsequent labour market outcomes using register data from Germany. I instrument for the conational share using hiring trends in the local labour market and find that a ten-percentage-point increase in the initial conational share lowers employment rates by 3.1 percentage points six or more years after the start of the first job, an effect not observed for non-conational immigrants, with no effect on wages conditional on employment. The employment effect appears to be due to the effect of differences in the composition of social networks induced by differences in the initial workplace on subsequent job search behaviour, although differential Germany-specific human capital acquisition cannot be entirely ruled out.

JEL-Codes: J610, J640, J310.

Keywords: employment, segregation, coworker networks, immigrant earnings dynamics.

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# 1 Introduction

A growing body of evidence has documented substantial segregation of workers across workplaces by country of origin in developed economies (Andersson et al., 2014; Åslund and Skans, 2010; Glitz, 2014; Hellerstein and Neumark, 2008). Immigrants are significantly more likely to work with other immigrants, and in particular immigrants from the same country of origin, than observable characteristics such as education, gender, or location would predict. Not only do immigrants have on average different coworkers when compared to natives, they also earn less than natives, in part because they are less likely to be employed (documented in Borjas, 1985; Chiswick, 1978; Lubotsky, 2007; Sarvimäki, 2011). A natural question is whether these two phenomena are related.

The relationship between the composition of an immigrant’s workplace and the immigrant’s labour market outcomes is, however, confounded by a number of factors. Immigrants may differentially select into first jobs with a higher or lower conational share based on unobserved characteristics related to future employability. Furthermore, the initial conational share is likely to be associated with other characteristics of the first job which might affect wages in that job, such as the presence of an immigrant manager (Åslund et al., 2014) or having received a referral (Dustmann et al., 2016). The true effect of the conational share on either contemporaneous or subsequent outcomes is therefore not identified by simple comparisons of immigrants who start in high- or low-conational share firms.

In this paper, I set out to provide credibly causal estimates of the effect of the conational share in the first job an immigrant holds in Germany on their subsequent labour market outcomes. To address the identification problem, I propose to instrument for the initial conational share using predicted hiring in the location and year where an immigrant is searching for a job, similar to the instrument proposed by Arellano-Bover (2020a) for the size of the firm where a worker finds her first job. Specifically, for a given immigrant, I calculate the expected share of conationals if the immigrant were randomly assigned a different job in their district that was filled by another immigrant in the same year. Conditional on fixed effects that capture selection into searching for a job in different labour markets based on time-varying nationality-specific factors and selection into districts based on the density of local ethnic networks, I provide evidence that the predicted conational share is quasi-randomly assigned.

The instrument relies on the idea that, conditional on when and where an immigrant decides to search for a job, there is some randomness in the set of firms closest to the immigrant that are looking to hire at that time. However, other firm characteristics may be correlated with the conational share, so my proposed instrument may predicted other firm characteristics too, violating the exclusion restriction. Furthermore, simply including supplementary characteristics of the firm where an immigrant holds her first job as additional controls in the structural equation would be invalid, since these characteristics are potentially outcomes of the instrument. To ensure that the exclusion restriction holds, I therefore adopt an idea used in judge leniency IV designs (Autor et al., 2015; Humphries et al., 2019) and use the same procedure as I used to calculated the predicted conational share to calculated a predicted version other firm-level characteristics.

The other predicted characteristics are then used as instruments for the other realised firm characteristics which are treated as endogenous variables in the structural equation, just like the conational share.

Implementing my empirical approach on a sample constructed from the German Sample of Integrated Employer-Employee Data (SIEED), I find that starting out in a firm with a higher conational share has a negative effect on an immigrant's probability of being employed in the longer term. A ten-percentage-point increase in the initial conational share reduces employment rates by 1.9 percentage points after two years, falling to 3.1 percentage points after six or more years. Importantly, the long-term employment effect is specific to the conational share and does not exist for immigrants who do not share the immigrant's nationality, suggesting that the underlying mechanism must be specific to the conational share. The estimates are robust to selective return migration, and descriptive evidence using survey data from the German Socio-Economic Panel (SOEP) suggests the effect is not due to an increase in self-employment. In contrast, there is no evidence of a wage effect for the conational share, while the share of other migrants is, if anything, positively associated with wages in the long-term.

Having established that the initial conational share has a negative longer-term effect on employment rates, I then review the evidence for different mechanisms that might explain this finding. The effect does not appear to be due to differences in productivity or wages in the first job, which might propagate to subsequent jobs via, e.g., job-ladder effects (Burdett and Mortensen, 1998). Rather, a higher initial conational share appears to worsen an immigrant's ability to search for jobs in the longer term. Immigrants with a higher initial conational share rely more on their former coworkers to find subsequent jobs, are more likely to transition from a job into unemployment rather than another job and, conditional on becoming unemployed, they tend to stay unemployed for longer. Again, these effects are not observed for the share of immigrants from other countries of origin. I interpret these effects as evidence that a higher initial conational share worsens an immigrant's job search capital in the future. Survey evidence from the SOEP suggests a higher conational share is not associated with worse German proficiency in the long-run. However, the conational share is negatively associated with having participated in formal job training in the longer run, so I cannot entirely rule out that part of the negative employment effect is due to differential host-country specific human capital accumulation, reducing immigrants' productivity in the longer run.

The first contribution of this paper is to provide plausibly causal estimates of the effect of workplace segregation on immigrants' outcomes. Previous work has shown that more segregated groups have worse labour market outcomes on average (Åslund and Skans, 2010; Glitz, 2014) and that higher conational shares in the first job are negatively associated with individual outcomes (Ansala et al., 2021). However, these associations are potentially confounded by the factors described above. This paper also relates to papers highlighting how working in a firm with more immigrants may benefit immigrants if their manager is also an immigrant (Åslund et al., 2014) or if it means receiving a referral when joining the firm (Dustmann et al., 2016). In contrast to those papers, I focus on the effect of the conational share *per se* and find it to be negative in the long run.

Second, this paper also contributes to a large literature studying how initial conditions upon arrival in a new country affect an immigrant’s career path. Typically, prior research has focused on the initial place of residence and the relationship between the size of an immigrant’s ethnic group in the initial location of residence and the immigrant’s subsequent labour market outcomes (Battisti et al., 2022; Beaman, 2012; Damm, 2009; Edin et al., 2003; Munshi, 2003). I extend this line of research from the neighbourhood to the firm. The switch of focus is novel; it is motivated in part by recent evidence that coworker networks are a more important determinant of an individual’s labour market outcomes than residential networks (Eliason et al., 2022).

Focusing on the firm, rather than the neighbourhood, is also motivated by the active literature on the role of firms for understanding earnings differences between immigrants and natives (Aydemir and Skuterud, 2008; Barth et al., 2012; Brinatti and Morales, 2021; Phan et al., 2022), to which I also contribute. Relative to these papers, which emphasise the role of sorting across high- or low-paying firms in determining immigrants’ contemporaneous wages, I show how a specific, time-varying characteristic of firms, namely the conational share at the time of first employment, has persistent long-term effects on an immigrant’s labour market outcomes and in particular employment. This is similar to the line of papers showing, for workers in general, that specific firm characteristics, and in particular the size of the firm, affect workers’ outcomes beyond the time of their employment in the firm (Arellano-Bover, 2020a,b).

The results presented here necessarily condition on an immigrant having found a first job. Nevertheless, the final contribution of the paper is to our understanding of immigrants’ labour market integration more broadly (see, e.g., Algan et al., 2010; Borjas, 1985; Chiswick, 1978; Dustmann and Görlach, 2015). In particular, the finding that a higher conational share lowers subsequent employment rates can help us understand the dynamic tradeoff that has been documented in relation to higher conational shares in the location of residence and immigrant earnings. Larger ethnic residential networks have been shown to increase earnings in the short-term but not in the long-term (Battisti et al., 2022). While a higher residential conational share may speed up job finding (e.g. Edin et al., 2003), I provide descriptive evidence, drawing on the SOEP, that immigrants who find work quicker do so in firms with a higher share of conational coworkers. This could act as a drag on their longer-term labour market performance and may afford immigrants who take longer to find work in a firm with fewer conationals a chance to catch up.

The paper proceeds as follows. In the following section I discuss the data used in the paper. In Section 3 I describe my empirical approach and challenges to identification. In Section 4 I present evidence on the relationship between initial workplace composition and subsequent employment rates and wages. In Section 5 I assess different possible mechanisms that could explain my result and relate my findings to the existing literature. Finally, Section 6 concludes.

## 2 Data

In the main analysis I use the Sample of Integrated Employer-Employee Data (SIEED), provided by the Institute for Employment Research (IAB) of the German Federal Employment Agency, which is described in detail in Schmidlein et al. (2020). The SIEED is constructed by first taking a 1.5 per cent sample of all firms making social security contributions during the period 1975–2018.<sup>1</sup> Second, the full employment biographies of all individuals ever employed by the sampled firms are then included in the dataset. I focus on immigrants whose first job was in one of the SIEED firms sampled at the first stage, for whom I observe the full set of coworkers in the first job, who were aged 15–64 at the time of this job, and who first appear in my dataset on or after 1 January 1991.<sup>2</sup> The administrative data only contains information on nationality, not migration status. Until a reform of the German nationality law in 2000, second-generation migrants frequently did not have German nationality. As a result, to avoid misidentifying immigrants, I exclude the major guest-worker countries, Turkey, Italy, and Greece from my sample, as the children of guest workers would be entering the labour market during my sample period. I also exclude individuals who ever report a foreign place of residence, to exclude commuters. The final sample includes around 39,000 individuals.

The employment biographies derived from the social security data only include employment in a job covered by the social security system. This means that work in self-employment or as a civil servant is not covered; breaks in employment biographies could therefore be indicative of unemployment, return migration, or employment in one of these categories. The data are reported as notifications, which record employment spells to the day. I transform the daily data into an annual panel, starting from the immigrant’s first year of social security-covered employment. In particular, I record the fraction of days worked in the calendar year, which I refer to as an individual employment rate, as well as the average daily wage earned in the course of the year, conditional on being employed at least one day. Firm-level variables are either calculated on 30 June, or on the day an individual started working in a firm, where relevant.

I report descriptive statistics in Table 1. All wage and earnings variables are deflated to 2010 values. Panel A presents time-varying statistics. The average employment rate in my sample, at 0.47, is lower than in the foreign born population as a whole, which averaged 0.64 during 2000–2018 (OECD, 2020). This reflects the fact that self-employment and return migration are not observed in the register data; individuals falling into either category are classified as unemployed. I will therefore present results that exclude individuals who drop out of employment permanently as a robustness check. Panel B presents time-invariant characteristics before the start of the first job. The sample contains a greater share of males than the immigrant population as a whole, reflecting the fact that labour force participation is higher among male immigrants than among female immigrants, while the educational distribution in the sample is similar to the that in

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<sup>1</sup>Formally, the SIEED samples establishments; an establishment corresponds to all production sites of a single employer in the same municipality operating in the same narrowly defined industry class. I follow convention when working with IAB data in referring to an establishment as a firm.

<sup>2</sup>The IAB data only cover East Germany from this date. I also exclude individuals who first appear in the dataset in East Germany on 1 January 1991, since these individuals were likely already working.



the wider immigrant population (OECD, 2020). Panel C presents characteristics of the first job or the firm where the first job is obtained. The first firm is on average large, with over 500 employees, however the distribution is highly skewed, and the median firm size is 62. Immigrants earn less on average in the first job (45 euros) than the median in the firm (64 euros).

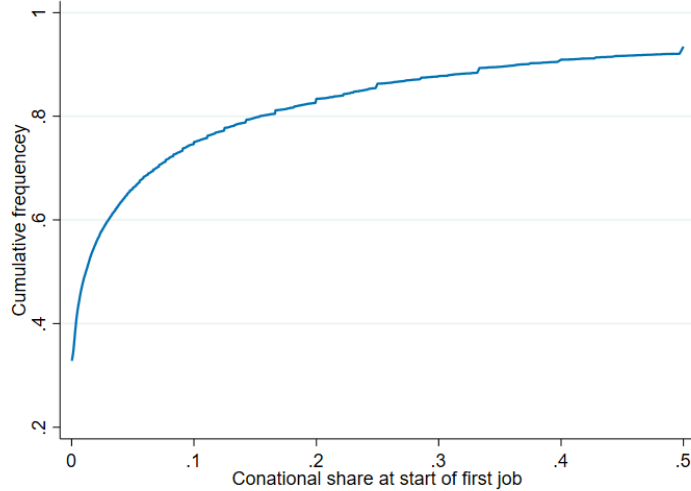
The average conational share in the first firm is 11 per cent and the average share of immigrants from other countries of origin is 22 per cent. In Figure 1 I further plot the cumulative distribution of the conational share in the first job, truncating the distribution at a conational share of 50 per cent. Just over 30 per cent of the sample do not have any conational coworkers in their first job, while around 10 per cent start in a workplace where the majority of their coworkers are conationals. Finally, I report the distribution of countries of origin in Table A.1. The largest groups of immigrants are from new members of the EU, with a fifth of the sample coming from Poland and Romania, with the next-largest group from the former Yugoslavia, making up around 12 per cent of the sample.

Table 1: Summary statistics

	Mean	SIEED St. dev.	N
Panel A			
Employment rate	0.47	0.45	626665
Annual wage earnings	11164.9	15871.5	626665
Avg. daily wage	60.7	49.6	373170
$\mathbf{1}(t \in [0, 2])$	0.19	0.39	626665
$\mathbf{1}(t \in [3, 5])$	0.19	0.39	626665
$\mathbf{1}(t \geq 6)$	0.63	0.48	626665
Panel B			
Woman	0.44	0.50	39069
low	0.58	0.49	39069
med	0.20	0.40	39069
high	0.22	0.41	39069
Age at first emp.	29.29	9.24	39069
Panel C			
$s_i^{own}$	0.11	0.21	39069
$s_i^{other}$	0.22	0.21	39069
Daily wage	45.1	39.6	39069
Apprentice	0.068	0.25	39069
Part-time	0.34	0.47	39069
Firm size	611.1	1900.8	39069
Median firm size	62	0	39069
Firm median wage	64.2	34.4	39069
Firm age	13.9	10.1	39069
Conat. manager	0.073	0.26	39069
Other mig. manager	0.11	0.32	39069

*Note:* Panel A reports time-varying summary statistics for the years since the first job, average earnings are conditional on being employed on June 30. Panel B reports summary statistics on individual characteristics at the start of the first job. Panel C reports summary statistics on the characteristics of the first job held after migration and the firm where the job was held. Wages and earnings are deflated and reported in 2010 Euros.

Figure 1: CDF of conational share in first job



*Notes:* Empirical CDF of the initial conational share in the first job held by an immigrant in my sample. The distribution is truncated at 50, for ease of representation.

In addition to the register data contained in the SIEED, I complement my analyses at certain points with survey data contained in the IAB-SOEP Migration Sample and linked to the social security data of the Institute for Employment Research (officially, the IAB-SOEP-MIG-ADIAB), which is described in detail in Brücker et al. (2013). The IAB-SOEP Migration Sample is an annual survey of individuals in Germany with a migration background (i.e. immigrants or descendants of immigrants), conducted as a supplement to the German Socio-Economic Panel (SOEP). Summary statistics on the 863 individuals in the linked IAB-SOEP data I use in supplementary analyses, who were born in a foreign country with a foreign nationality and who arrived in Germany between the ages of 15 and 64, are contained in Table A.2, while the distribution of the initial conational share in the IAB-SOEP is in Figure A.1 and the distribution of nationalities is in Table A.3.

### 3 Empirical approach

#### 3.1 Overview

To estimate the effect of the initial conational coworker share on immigrants' subsequent labour market outcomes, I model outcomes of interest  $t$  years after the start of  $i$ 's first job,  $Y_{it}$ , as a function of the initial conational share  $s_i^{own}$ . In general, the outcome can be assumed to follow some nonparametric time trend,  $f_2(t)$ , and the effect of interest,  $f_1(t)$ , likewise may be non-constant over time. Furthermore, other time-varying and invariant individual factors  $X_{it}$  may affect the outcome of interest. Finally, more aggregate fixed characteristics, such as cohort effects, if the "quality" of immigrant is changing over time, nationality, or location of arrival within Germany, measured as fixed effects  $\delta_j$  may affect the outcome. Putting it all together,

we have

$$Y_{it} = f_1(t) \times s_i^{own} + f_2(t) + \Gamma X_{it} + \sum_j \delta_j + \epsilon_{it}. \quad (1)$$

To make the estimation problem more tractable, I adopt a semi-flexible approach to modelling the functions  $f_1(t)$  and  $f_2(t)$ . Ideally, I would like to model each as a set of indicator variables for all values that  $t$  takes on. However, I would then have insufficient power to identify the large set of effects of interest. I therefore group years together and instead model both functions as a set of indicator variables for being within 0-2 years of the first job, 3-5 years of the first job, or more than 6 years of the first job. Within the set of control variables  $X_{it}$ , I will pay special attention to the share of immigrants from other countries of origin in the first job,  $s_i^{other}$ , whose effect I will allow to vary over time, just as the effect of the conational share does. The initial share of other immigrants is of special interest, as compared to other characteristics of the firm where the first job is held. A large part of immigrant segregation is due to the excess tendency of immigrants to work with their conationals specifically. It is therefore important to understand whether the effect of exposure to conationals is different to the effect of exposure to immigrants in general. The final estimating equation is thus

$$Y_{it} = \sum_{g \in \{own, other\}} \beta_1^g s_i^g \times \mathbf{1}(t \in [0, 2]) + \beta_2^g s_i^g \times \mathbf{1}(t \in [3, 5]) + \beta_3^g s_i^g \times \mathbf{1}(t \geq 6) + \mathbf{1}(t \in [0, 2]) + \mathbf{1}(t \in [3, 5]) + \mathbf{1}(t \geq 6) + \Gamma X_{it} + \sum_j \delta_j + \epsilon_{it}. \quad (2)$$

The share variables,  $s_i^g \in [0, 1]$ , implying that the reported coefficient  $\beta_\tau^g$ ,  $\tau \in \{1, 2, 3\}$ , measures the effect of going from a firm with no coworkers of type  $g$  to an equivalent firm made up entirely of coworkers of type  $g$ . In the text, unless otherwise stated, I will scale this coefficient and discuss the effect of a ten-percentage-point increase, or approximately half a standard deviation, in the share of workers of type  $g$ . In all cases, the set of control variables will include basic demographic characteristics, gender and a quadratic in age, and pre-employment characteristics, educational attainment at the start of the first job and age at the start of the first job.

## 3.2 Identification

In practice, not all relevant elements of  $X_{it}$  are available. The SIEED does not contain information on most of a worker's relevant pre-migration characteristics, such as German proficiency, or how they found their first job and whether they received a referral. Furthermore, individual preferences, such as a taste for working with conationals, or fixed individual characteristics, such as employability in Germany, are not observable. As a result, OLS estimates of Equation (2) are likely to yield biased estimates of  $\beta_\tau^g$ . I therefore adopt an instrumental variables (IV) approach to identifying the effect of the initial conational share on subsequent outcomes. The proposed instrument uses variation across districts (*Kreise*) within the same labour market in the hiring patterns of firms for a given year and nationality. Formally, the instrument is defined as

follows:

$$z_i^{own} = \frac{\sum_{j \neq i} s_{f(j)}^{nat.(i)} \mathbf{1}(d_j = d_{0i}, t_j = t_{0i}, nat.(j) = mig.)}{\sum_{j \neq i} \mathbf{1}(d_j = d_{0i}, t_j = t_{0i}, nat.(j) = mig.)} \quad (3)$$

The instrument for individual  $i$  is the average share of coworkers with  $i$ 's nationality among other migrants  $j \neq i$  hired in the same district as  $i$ ,  $d_{0i}$  and the same year,  $t_{0i}$ .<sup>3</sup> The instrument is therefore a leave-out-mean and has the same structure as the instrument proposed by Arellano-Bover (2020a) for the size of the establishment where Spanish school-leavers find their first job. To avoid contamination by hires throughout the year, the share of conationals  $s_{f(j)}^{nat.(i)}$  is measured on January 1 and the instrument is constructed using hires during the calendar year. The instrument can be interpreted as the expected conational share if an immigrant were randomly assigned to a position filled in the same year by another immigrant; throughout the paper I refer to it as the predicted conational share.

The predicted conational share in a district is correlated with other characteristics of the district (see Figure A.2). For example, a higher predicted conational share is marginally positively correlated with the conational share in the district's labour force, which would indicate a larger network for the immigrant to draw on when searching for a job (see, e.g., Munshi, 2003). To ensure that other district characteristics do not confound the effect of the conational share on subsequent outcomes, my main specification will therefore include fixed effects for labour market<sup>4</sup> by nationality by year of first job and for district (*Kreis*) by nationality. The quasi-random assignment component of the IV identifying assumptions therefore requires that, for individuals of the same nationality, searching for a job in the same labour market in the same year, the deviation from its average across the sample years of the conational coworker share predicted by firm hiring in the district the immigrant is searching in is uncorrelated with individual characteristics of the immigrant that predict longer-term labour market outcomes. Put differently, workers of a given nationality, searching in the same labour market and year should not systematically sort into districts within that labour market where firms that employ relatively many of their conationals are doing a disproportionately large or small share of hiring that year, relative to the long-term average of the district.

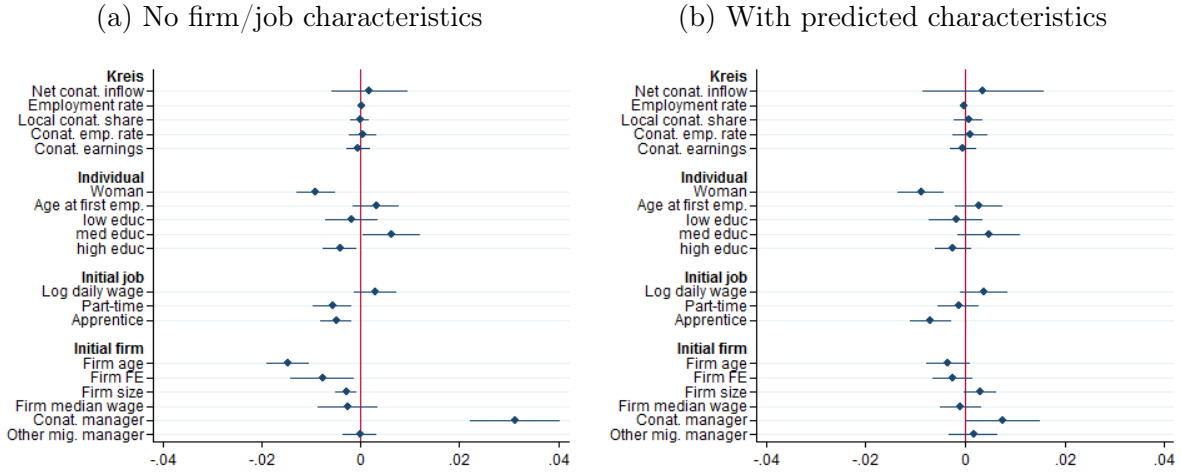
To assess the validity of the instrument, I report results in Figure 2a from a series of cross-sectional regressions where I regress a set of standardised variables measuring characteristics of the district where the first job is found, individual characteristics (measured at the start of the first job), and characteristics of the first job and first firm on the predicted conational share and the set of fixed effects described above. If conditional random assignment holds, the instrument should not be associated with other characteristics of the district at the time of finding the job, or of the individual, conditional on the included fixed effects. This is largely what we observe. The instrument is conditionally uncorrelated with other local characteristics such as the inflow

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<sup>3</sup>Note the instrument is constructed using immigrant hires across the entire set of SIEED firms, not only immigrants hired for the first time. The estimation sample is therefore very small relative to the sample used to construct the instrument, so the two samples can be considered effectively independent when conducting inference.

<sup>4</sup>Labour markets are defined by commuter flows, see (Kropp and Schwengler, 2011); there are 50 labour markets in Germany.

Figure 2: Instrument validity



*Notes:* Effect of predicted conational share on other characteristics. Each association is estimated separately; the dependent variable in each specification has been standardised to have mean 0 and standard deviation 1, while the predicted share is rescaled to lie on  $[0,100]$ . All specifications include labour market  $\times$  nationality  $\times$  entry year and district  $\times$  nationality fixed effects. Standard errors are clustered by entry district.

of conationals into the district in the same year, local employment rates, or conationals' share of either the local population or the local labour force. The individual characteristics are also largely uncorrelated with the predicted conational share, with the exception of gender. A ten-percentage-point increase in the predicted conational share reduces the probability an immigrant is a woman by around 0.1 of a standard deviation. In addition to controlling for the individual characteristics, I will also assess whether or not effects are heterogeneous by gender or initial education level.

While the results reported in Figure 2a suggest that the the instrument is conditionally randomly assigned, they also show that the exclusion restriction does not hold. Specifically, the predicted conational share predicts characteristics of the firm where the first job is held other than the realised conational share. Some of these characteristics may have a direct effect on immigrants' subsequent labour market outcomes, such as the size of the firm where the first job is held (Arellano-Bover, 2020a) or the presence of a conational manager at the firm (Åslund et al., 2014), in which case the exclusion restriction will be violated. To address this problem, it is possible to use the same leave-out-mean procedure to calculate a predicted version of any initial job or firm characteristic. The realised job and firm characteristics can then be included as controls in Equation (2), the structural equation, and instrumented for, using their predicted versions, in the IV estimation.<sup>5</sup> In Figure 2b, I include predicted part-time status, firm age, presence of a conational manager, presence of an immigrant manager from another country, and the log of predicted firm size and the predicted median wage in the firm, all calculated using the

<sup>5</sup>Note that it would not be correct to include job or firm characteristics directly as controls in Equation (2) without instrumenting for them. Since these characteristics are outcomes of the proposed instrument, they would be bad controls in the reduced-form equation, biasing the two-stage least squares estimate of the effect of the conational share.

same leave-out-mean procedure, as additional controls in my cross-sectional regressions. When including the predicted characteristics as additional controls, the predicted conational share is no longer systematically associated with the job and firm characteristics. I report the same tests of instrument validity for the predicted share of workers from other countries of origin in Figure A.3. Again, there is some evidence, albeit weaker, of the exclusion failing to hold when other predicted firm characteristics are not controlled for.

Turning now to assessing the relevance of the proposed instrument, I report the results of cross-sectional regressions of the realised conational share in the first job on the predicted conational share in Table 2. In column one I report the bivariate relationship between the the predicted and actual conational share; the raw correlation is equal to 0.44 Moving through columns 2–7, I progressively include more restrictive sets of fixed effects and, finally, controls for other predicted job and firm characteristics. Throughout,  $R^2$  rises from 0.19 to 0.65, however the instrument is highly significant and continues to predict the actual conational share almost one-to-one. I repeat the same set of regressions for the share of immigrants coming from other countries among the set of coworkers in the first job and report the results in Table A.4. While the relationship is a little weaker, the predicted share of other immigrants is nevertheless strongly predictive of the actual share of other immigrants.

Table 2: First stage effect of predicted conational share on realised share.

	(1)	(2)	(3)	(4)	(5)	(6)
$z_i^{own}$	1.31** (0.078)	1.30** (0.079)	1.26** (0.086)	1.10** (0.077)	1.15** (0.12)	0.99** (0.070)
$z_i^{other}$		-0.0016 (0.027)	0.024 (0.022)	0.022 (0.021)	0.024 (0.032)	0.037+ (0.022)
Pred. charact.	No	No	No	No	No	Yes
N	39069	39069	39069	39069	37302	39069
$R^2$	0.19	0.19	0.52	0.60	0.65	0.60
FE	-	-	LxNxY, D	LxNxY, DxN	LxNxY, DxN, DxY	LxNxY, DxN

*Note:* Static first-stage relationship between predicted conational share,  $z_i^{own}$ , and the realised conational share in the first job. Included predicted characteristics are part-time status, firm age, conational manager, other migrant manager, log predicted firm size and log predicted median wage. L = labour market, N = nationality, D = district, Y = year of first job. Standard errors clustered by district +  $p < 0.1$ , \*  $p < 0.05$ , \*\*  $p < 0.01$

Finally, when estimating the dynamic effect of the conational share, the predicted conational share and other immigrant share will be interacted with the same set of time-since-migration dummies as the actual shares, as in Equation (2). Turning from estimation to inference, I report standard errors clustered at the district level. Strictly speaking, the value of the instrument varies for each individual, however the firm-level hiring shocks from which the instrument is constructed are common to immigrants finding a job in the same labour market in the same year, suggesting that the district-year is the level at which treatment is assigned and standard errors should be clustered (Abadie et al., 2017). However, firm-level labour demand shocks may be somewhat persistent over time, leading to some serial correlation in the instrument, which leads

me to cluster at the district level.

## 4 The effect of the initial conational share

### 4.1 Employment rates

The main outcome of interest is individual employment rates. An individual's employment rate is defined as the fraction of days they are employed in a job covered by social security in a calendar year. I first report estimates of the reduced form effect of the predicted conational share in Panel A of Table 3. All specifications include a labour market by nationality by year of first job fixed effect, while in columns 2–4 I sequentially include district, district by nationality, and district by nationality and district by first year fixed effects. The pattern of effects is relatively consistent; the predicted conational share has a negative effect on subsequent employment rates and this effect becomes more negative over time. However, the effect attenuates somewhat as I include more detailed fixed effects. For example, six or more years after the first job, the a ten-percentage-point increase in the conational share lowers employment by 3.3 percentage points in the basic specification, and 2.3 percentage points when the full set of fixed effect is included. In columns 5–7 I also include other predicted characteristics of the firm and job. Now, including more-detailed fixed effects does not affect the estimated reduced form effect much, suggesting the predicted characteristics, in addition to ensuring the exclusion restriction holds, may also capture some residual selection into districts that the fixed effects capture when the predicted characteristics are not included. The reduced form effect of the predicted share of immigrants from other countries of origin follows a different pattern. Across the different specifications, the short-term effect is negative and generally smaller than the effect of the predicted conational share, while the long-term effect is statistically and economically insignificant.

Turning to the effect of the realised conational share, I report OLS estimates of the association between the realised conational share and subsequent employment rates in column one of Panel B, including realised firm characteristics as controls as well as labour market by nationality by first year and nationality by district fixed effects. The same time pattern is observed as for the predicted conational share. A ten-percentage-point increase in the realised conational share lowers employment rates by 0.9 percentage points in the short-term and by 2 percentage points in the long-term.

In columns 2–7 of Panel B, I report the 2SLS estimates. In columns 2–4 I do not include other firm characteristics as controls, while in columns 5–7 I include other firm characteristics and instrument for them using predicted characteristics. The 2SLS estimates are typically larger in magnitude, i.e. more negative, than the OLS estimates. This could be due to the fact that finding a job in a firm with a higher conational share may be a proxy for receiving a referral, which raises subsequent employment rates (Dustmann et al., 2016) or having an immigrant manager, which lowers separations (Åslund et al., 2014), both of which would bias the OLS estimates upwards. The pattern of 2SLS estimates across specifications is similar to the pattern of reduced-form estimates, in Panel A. Including more-detailed fixed effects reduces the mag-



nitude of the coefficients somewhat, however the long-term effect remains negative, significant, and larger in magnitude than the short-term effect. Including controls for other firm characteristics increases the estimated magnitude slightly, however the joint-first stage Kleibergen-Papp F-statistic is smaller, as the instrument set is larger and not all realised characteristics are equally well predicted by the equivalent predicted characteristic. Column 5 is my preferred specification, repeating the OLS specification and instrumenting for all firm characteristics, ensuring the exclusion restriction holds. According to these estimates, a ten-percentage-point increase in the realised conational share lowers short-term employment rates by 1.9 percentage points and long-term employment rates by 3.1 percentage points. Finally, in column 7 I include a fixed effect for the year in which the outcome is observed, in case different cohorts are exposed to the national business cycle at different points in time since their arrival. The estimated effect is almost identical to my preferred specification.

To put the magnitude of the long-term employment effect into context, Glitz (2014) finds that the average employed immigrant in Germany in 2008 had 18 percentage points more conational coworkers than would be expected under a random allocation of workers, or 13 percentage points after partialling out the effects of region of residence, gender, education, and industry. The employment rate of the foreign-born in Germany at the time was 62.9 per cent, 8.7 percentage points lower than the employment rate of the native-born (OECD, 2020). Scaling the long-term effect of the conational share in my preferred specification by average segregation translates to an employment rate that is  $0.31 \times 18 = 5.6$  percentage points lower, or 4 percentage points if observable characteristics are partialled out of the measure of segregation. The magnitude of the long-term association between the initial conational share and employment is therefore large relative to the difference in employment rates between immigrants and natives in Germany.

Finally, the estimated effect of the share of immigrants coming from other countries of origin is negative in the short-run, implying a 1.7-percentage-point decrease in employment rates for a ten-percentage-point increase in the conational share, but is zero in the longer-run. This difference between the effect of conationals and the effect of other immigrants constitutes a novel finding. Furthermore, it will be important to bear this difference in mind when evaluating potential mechanisms, since it implies that any mechanism that explains the effect needs to be specific to the conational share, and cannot apply to immigrants in general.

## 4.2 Robustness

As noted in Section 2, the SIEED does not record return migrants or the self-employed. As a result, the negative employment effect of the conational share could at least in part be due to immigrants leaving the country or shifting to self-employment.<sup>6</sup> In column 1 of Table 4, I repeat my main IV specification using a dummy for having dropped out of employment permanently, according to the SIEED, as an outcome. I find that a ten-percentage-point increase in the initial

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<sup>6</sup>Note, however, that return migration and, to a lesser extent, self-employment are also indicative or reduced success in the labour market for immigrants. As such, the negative effect of the conational share on subsequent SIEED employment is still a measure of reduced labour market success, even if part of the effect were to be interpreted as increased return migration or self-employment.

Table 3: Individual annual employment rates

	OLS		2SLS				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
<i>Panel A: Reduced form</i>							
$\mathbf{1}(t \in [0, 2]) \times z_i^{own}$		-0.17** (0.060)	-0.14* (0.055)	-0.075 (0.082)	-0.16* (0.072)	-0.16 (0.10)	-0.15* (0.071)
$\mathbf{1}(t \in [3, 5]) \times z_i^{own}$		-0.30** (0.054)	-0.27** (0.059)	-0.20* (0.088)	-0.28** (0.074)	-0.28* (0.12)	-0.28** (0.074)
$\mathbf{1}(t \geq 6) \times z_i^{own}$		-0.33** (0.073)	-0.30** (0.073)	-0.23* (0.099)	-0.31** (0.094)	-0.32* (0.13)	-0.31** (0.091)
$\mathbf{1}(t \in [0, 2]) \times z_i^{other}$		-0.11* (0.044)	-0.11* (0.046)	-0.17** (0.058)	-0.12* (0.051)	-0.17* (0.070)	-0.033 (0.048)
$\mathbf{1}(t \in [3, 5]) \times z_i^{other}$		-0.054 (0.038)	-0.053 (0.041)	-0.12* (0.058)	-0.068 (0.049)	-0.11 (0.070)	-0.047 (0.048)
$\mathbf{1}(t \geq 6) \times z_i^{other}$		-0.012 (0.034)	-0.011 (0.037)	-0.083 <sup>+</sup> (0.050)	-0.026 (0.044)	-0.079 (0.060)	-0.079 <sup>+</sup> (0.043)
<i>Panel B: OLS and 2SLS estimates</i>							
$\mathbf{1}(t \in [0, 2]) \times s_i^{own}$	-0.088** (0.023)	-0.13** (0.046)	-0.13** (0.047)	-0.055 (0.067)	-0.19* (0.086)	-0.23 (0.18)	-0.19* (0.087)
$\mathbf{1}(t \in [3, 5]) \times s_i^{own}$	-0.16** (0.025)	-0.23** (0.041)	-0.23** (0.050)	-0.15* (0.071)	-0.28** (0.086)	-0.33 <sup>+</sup> (0.18)	-0.29** (0.086)
$\mathbf{1}(t \geq 6) \times s_i^{own}$	-0.20** (0.024)	-0.25** (0.050)	-0.26** (0.060)	-0.18* (0.079)	-0.31** (0.10)	-0.36 <sup>+</sup> (0.19)	-0.30** (0.10)
$\mathbf{1}(t \in [0, 2]) \times s_i^{other}$	-0.13** (0.017)	-0.17* (0.072)	-0.16* (0.078)	-0.28** (0.096)	-0.17* (0.085)	-0.21 <sup>+</sup> (0.13)	-0.043 (0.083)
$\mathbf{1}(t \in [3, 5]) \times s_i^{other}$	-0.065** (0.018)	-0.092 (0.067)	-0.088 (0.076)	-0.20* (0.096)	-0.093 (0.086)	-0.14 (0.13)	-0.065 (0.084)
$\mathbf{1}(t \geq 6) \times s_i^{other}$	-0.037 <sup>+</sup> (0.019)	-0.032 (0.064)	-0.029 (0.071)	-0.15 <sup>+</sup> (0.086)	-0.035 (0.082)	-0.088 (0.12)	-0.12 (0.081)
Firm characteristics	Yes	No	No	No	Yes	Yes	Yes
Observations	501605	501605	501605	501605	501605	501605	501605
Individuals	39069	39069	39069	39069	39069	39069	39069
KP F-statistic		55.5	46.2	36.1	14.0	6.3	14.0
FE	NxD	D	NxD	NxD, DxY	NxD	NxD, DxY	NxD, oY

*Notes:* Each coefficient measures the effect of a one-percentage-point increase the share of coworkers on a given type on subsequent employment rates, measured in percentage points. Firm characteristics are part-time status, firm age, conational manager, other migrant manager, firm size, and median wage. L = labour market, N = nationality, D = district, Y = year of first job, oY = year of observation; all specifications include a LxNxD fixed effect. Standard errors are clustered by district. +  $p < .1$ , \*  $p < .05$ , \*\*  $p < .01$

conational share does indeed increase the probability of dropping out of formal employment altogether, that this effect is increasing over time, and that there is no such effect for the other immigrant share. In column 2 I therefore restrict my sample to those individuals who have not yet dropped out altogether, i.e. those either working, or currently unemployed but who will be observed returning to formal employment in the future. These estimates cannot be interpreted causally, since I condition on an outcome of the variable of interest. Nevertheless, the initial conational share remains negatively associated with subsequent employment rates, and the effect becomes more negative over time, while the negative short-term association with the other immigrant share is again transient.

Another perspective on the relationship between the initial conational share, return migration, and self-employment is provided by the IAB-SOEP data. There is no scope for return migration in these data, since they are constructed by surveying immigrants still in Germany in 2013 and 2014 and then matching their survey responses retrospectively to their social security data. However, the dataset is too small to use the estimation strategy described in Section 3.2, which relies on a relatively detailed set of fixed effects. I therefore estimate descriptive regressions on the IAB-SOEP data using OLS, where, in addition to the controls included in the IV specification, I further control for pre-migration characteristics recorded in the SOEP and otherwise unavailable in the social security data, such as German proficiency, having contacts in Germany prior to migrating, and indicators for different methods of finding the first job in Germany, e.g. through contacts or through the public employment agency.

I report estimates of the dynamic association of the conational and other migration shares with employment rates in column 3 of Table 4. The results are not directly comparable to the IV estimates using the SIEED, in light of the differences in sample construction and identifying variation. However, even in a sample where all individuals are known to still be in Germany at the end of the sample period, the initial conational share is still negatively associated with subsequent employment rates and the association becomes more negative over time; the other immigrant share is not significantly associated with subsequent employment at any time horizon.<sup>7</sup> The SOEP also contains information on employment as a civil servant or in self-employment for 2013 and 2014, the categories of employment not covered in the SIEED. I use an indicator for these types of employment as the outcome in columns 4 and 5. The association of both share variables with employment in the civil service, in column 5 is quite precisely estimated to be zero. The estimated association with self-employment is more noisy, however the coefficients for the conational share are positive and, for 3–5 years after entering employment, significantly so. One therefore cannot rule out that at least part of the negative employment effect estimated on the SIEED is due to an increase in self-employment.<sup>8</sup>

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<sup>7</sup>If the conational share in the first job has a positive effect on return migration, as the SIEED estimates seem to suggest, then the effect of the conational share on subsequent employment may be underestimated in a sample that, like the IAB-SOEP Migrant Sample, conditions on not having return migrated, as I discuss in Appendix B.

<sup>8</sup>Andersson (2021) finds that refugees' self-employment is positively affected by the share of self-employed coethnics in the municipality of entry, but not by the share of co-ethnics per se. There may therefore be no strong reason *a priori* to presume that a higher share of conationals in the first job in

Table 4: Other measures of employment and labour force participation

	(1)	(2)	(3)	(4)	(5)
	Dropout	Emp.   no dropout	Employed	Self-employed	Civil servant
$\mathbf{1}(t \in [0, 2]) \times s_i^{own}$	0.045 (0.083)	-0.26** (0.083)	-0.0080 (0.058)	0.15 (0.16)	-0.0030 (0.013)
$\mathbf{1}(t \in [3, 5]) \times s_i^{own}$	0.21* (0.096)	-0.28** (0.083)	-0.071 (0.067)	0.38* (0.18)	-0.030 (0.019)
$\mathbf{1}(t \geq 6) \times s_i^{own}$	0.24* (0.11)	-0.30** (0.10)	-0.14* (0.070)	0.076 (0.061)	0.0070 (0.0077)
$\mathbf{1}(t \in [0, 2]) \times s_i^{other}$	0.048 (0.073)	-0.24** (0.061)	-0.030 (0.045)	0.021 (0.035)	0.014 (0.014)
$\mathbf{1}(t \in [3, 5]) \times s_i^{other}$	-0.041 (0.073)	-0.12+ (0.067)	-0.053 (0.059)	-0.055 (0.041)	-0.012 (0.013)
$\mathbf{1}(t \geq 6) \times s_i^{other}$	-0.070 (0.077)	-0.040 (0.067)	-0.060 (0.064)	-0.0053 (0.032)	-0.0074 (0.0060)
Observations	501605	368720	10061	1506	1506
Individuals	39069	39069	863	849	849
KP F-statistic	14.04	13.11	–	–	–
Source	SIEED	SIEED	IAB-SOEP	SOEP	SOEP

*Notes:* Each coefficient measures the effect of a one-percentage-point increase the share of coworkers. Outcomes measured in the SIEED are estimated using 2SLS following Equation (2). Outcomes measured in IAB-SOEP sample are estimated using OLS, including additional controls for measured characteristics. Standard errors are clustered by district in the SIEED and by individual in the IAB-SOEP. +  $p < .1$ , \*  $p < .05$ , \*\*  $p < .01$

Having established that the estimated effect of the conational share on employment is unlikely to be fully explained by either return migration or self-employment, I assess whether there is any heterogeneity in the effect by other characteristics of the individual or firm and report the estimates in Table A.5. In order to have sufficient power to test for heterogeneous effects, I abstract from the dynamic effect of the conational share and estimate a cross-sectional regression, where the dependent variable is the average employment rate over the first eight years since the start of the first job. The baseline effect is reported in column 1; consistent with the dynamic specification, a ten-percentage-point increase in the conational share lowers employment rates by 3.3 percentage points, while the other migrant share has no effect. In column 2 I confirm that these effects are homogeneous by gender. In column 3 I show that the effect is larger for less-educated immigrants; for highly educated immigrants, a ten-percentage-point increase in the conational share lower employment by a statistically insignificant 2 percentage points. Finally, in column 4 I allow for heterogeneity by the size of the first firm. Since the size of the first firm is potentially an outcome of the instrument, these estimates cannot be interpreted causally. Nevertheless, the estimated coefficient on the conational share is larger in firms with fewer than 100 employees, consistent with the conational share mattering due to actual contact with one's coworkers. In larger firms, the conational share will less-accurately capture the degree of contact one has with conational coworkers in the workplace.

I also consider whether the conational share is proxying for other characteristics of coworkers by including other measures of average network quality in the local area or in the first firm in the cross-sectional regression. Specifically, I consider the employment rate in the district in the year of the first job, the share of conationals in the district population in the year of the first job, and the average employment rate of one's coworkers according to the SIEED over the five years prior to the start of the first job. I report the results in Table A.6, where the included measures are first standardised to have mean zero and standard deviation one. In columns 2–4 I include district or coworker characteristics as controls in the same cross-sectional IV specification as previously. Including these controls does not materially alter the effect of the conational share, even if it becomes less strongly significant when conditioning on the conational share in the district population, since these two variables are strongly correlated.

In columns 5–7 I also interact the included extra controls with the conational and other migrant share in the first job. The only extra characteristic with a significant main effect in these specifications is the employment rate of an immigrant's coworkers before the immigrant joins the firm, in column seven, however the main effects of the coworker shares are not materially affected and the interaction terms are small and insignificant. The largest change in the effect of the coworker shares in these specifications occurs when interacting the conational share at the firm with the district-level conational share, in column six. In this case, the main effect of a ten-percentage-point increase in the initial conational share is now a statistically insignificant 2.5-percentage-point decrease in employment rates. Furthermore, interestingly, the effect of the conational share in the first job becomes less negative when the local conational share is lower,

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formal employment might have an effect on subsequent self-employment, since conationals in the first job are themselves not in self-employment, at least initially.

although the interaction term is not significant by itself. This might be because conationals encountered on the job are better-integrated when the local conational share is lower, or because a newly-arrived immigrant has more opportunities to interact with natives outside of work when the local conational share is lower, undoing the effect of more-intensive interactions with conationals in the workplace. The general picture to emerge from Table A.6, however, is that the conational share in the first job is not proxying in a systematic way for some other characteristic of one’s initial set of coworkers.

Finally, I also assess the robustness of various assumptions I make about the functional form, embedded in Equation (2). First, the effect may be non-monotonic in the conational share (c.f. Ansala et al., 2021). In Figure A.4 I plot the average employment rate for different categories of the initial conational share, conditional on included controls. All averages are expressed as deviations from the employment rate of individuals whose initial conational share is less than 5 per cent in their first two years of employment.<sup>9</sup> The association between the initial conational share and long-term employment rates does appear to be monotone.

Second, the grouping of time dummies in Equation (2) may be overly restrictive. I estimate a specification by OLS where I allow the effect of both group shares to vary for each year since the start of the job. The estimated coefficients are reported in Figure A.5. The time pattern of effects is similar to what I observe with the simpler specification, although there is a clear drop-off in the association between the initial conational share and employment rates between years zero and one that is obscured by the grouping of time dummies.

### 4.3 Wage earnings

In the aggregate, immigrants not only have lower employment rates than native, but also have lower wages conditional on employment (Algan et al., 2010). I therefore repeat my main specification using different measures of wages as outcomes, conditional on employment. In the light of the effect of the conational share on subsequent employment rates, documented in Table 3, conditioning the analysis on employment, while necessary, is problematic, since it implies that the sample is endogenously selected. The association between the initial conational share and wages should therefore not be interpreted causally.

The results of the wage analysis are reported in Table 5. The social security data only include daily wages, rather than hourly wages, and an indicator for part-time status. In column 1 I therefore estimate the association between the initial conational share and average daily earnings, defined, for individuals who work at least one day during the year, as total earnings subject to social security in a year divided by total number of days worked, deflated to 2010 values. In columns 2 and 3 I condition on working either full-time or part-time on June 30, to account for the effect of hours worked in a day, and take as my measure of wages wage earnings on

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<sup>9</sup>These averages are estimated by replacing the interactions of the conational share with years since migration in Equation (2) with a full set of interactions between the years since migration and a set of dummies for the base year immigrant share taking values from [0, 5), [5, 10), [10, 50), [50, 90), and [90, 100]; individuals with a conational share in the 0–5 per cent range in their first two years of employment are the omitted category. The specification is estimated by OLS.

June 30. There is a short-term positive association between the initial conational share and the daily wage, particularly for full-time workers, however there is no statistically or economically significant longer-term association, conditional on being employed. The share of immigrants coming from other countries appears, if anything, negatively associated with wages conditional on employment and the association appears to grow somewhat over time, particularly for part-time workers. However, the magnitudes are small—a ten-percentage-point increase in the other migrant share is associated with 0.6 percent lower daily wages—and, pooling full-time and part-time workers, is statistically insignificant in both the short- and the long-run.

In columns 4–6 I repeat the same specifications, this time instrumenting for the conational share, the other immigrant share, as well as the same set of other firm characteristics as in the employment analysis (firm age, median wage, size, manager nationality) and part-time status in the first job. Now there is no significant association between the initial conational share and average daily wages at any time horizon. Only when conditioning on working part-time do I find a statistically significant association between the initial conational share and wages and only in the long term. The association is negative and the magnitude is quite large, A ten-percentage-point increase in the initial conational share is associated with an 11 percent decrease in the daily wage, significant at the five per cent level. However, a relatively weak first stage means this isolated result should be interpreted with caution. On the other hand, the share of immigrants from other countries of origin appears to be negatively associated with daily wages in the short term and positively associated with daily wages in the long term. Here a ten-percentage-point increase in the other migrant share is associated with a marginally significant 2.8 percent increase in wages for full-time workers and a 5.8 percent increase in wages for part-time workers.

As noted previously, conditioning the analysis on any employment creates a selection bias. Individuals who are employed, whether full-time or part-time, in spite of having a high conational share in their first job are potentially positively selected on unobserved employability relative to other immigrants, introducing a conditional-on-positive selection bias (Angrist and Pischke, 2009). This kind of selection would likely bias the estimated association between the initial conational share and potential subsequent earnings upward relative to the true association in the full, unobservable, population. As such, while the associations presented in Table 5 provide *prima facie* evidence against a wage effect of the initial conational share, it is ultimately not possible to conclude whether the true causal effect of the initial conational share on wages is zero, as these estimates suggest, or negative but biased toward zero when conditioning on individuals being employed.

To conclude, the pattern of associations of the share variables with long-term wages conditional on employment—a null or negative association for the conational share, and if anything a positive association for the share of immigrants from other countries—does not undo the reduction in total wage earnings implied the negative employment effect of the conational share. Furthermore, the finding of a clear negative effect of the starting conational share on long-term employment, contrasting with limited evidence of a wage effect is consistent with evidence that the total earnings gap between immigrants and natives is mostly due to differences in employment, not wages conditional on employment (Sarvimäki, 2011).

Table 5: Relation between initial workplace composition and log wages

	OLS			2SLS		
	(1)	(2)	(3)	(4)	(5)	(6)
$\mathbf{1}(t \in [0, 2]) \times s_i^{own}$	0.22** (0.041)	0.14** (0.042)	0.089 (0.073)	-0.29 (0.20)	-0.22 (0.18)	-0.22 (0.45)
$\mathbf{1}(t \in [3, 5]) \times s_i^{own}$	0.11** (0.030)	0.058+ (0.034)	0.093+ (0.049)	-0.11 (0.17)	0.043 (0.17)	-0.39 (0.45)
$\mathbf{1}(t \geq 6) \times s_i^{own}$	0.010 (0.040)	-0.015 (0.044)	0.025 (0.056)	-0.30 (0.27)	0.052 (0.23)	-1.08* (0.43)
$\mathbf{1}(t \in [0, 2]) \times s_i^{other}$	-0.027 (0.046)	-0.085* (0.039)	-0.0038 (0.085)	-0.30* (0.14)	-0.41** (0.13)	0.027 (0.20)
$\mathbf{1}(t \in [3, 5]) \times s_i^{other}$	-0.10** (0.037)	-0.15** (0.036)	-0.11* (0.056)	-0.089 (0.14)	-0.28* (0.12)	0.17 (0.18)
$\mathbf{1}(t \geq 6) \times s_i^{other}$	-0.062 (0.047)	-0.066 (0.044)	-0.16* (0.064)	0.51** (0.17)	0.27+ (0.14)	0.56** (0.19)
Observations	316315	216346	100352	316315	216346	100352
Individuals	39068	33736	20944	39068	33736	20944
KP F-statistic				14.4	9.7	4.2
Subsample	all	FT	PT	all	FT	PT

*Note:* Columns 1–3 report OLS estimates of relationship between initial conational share and log wages, conditional on employment, columns 4–6 report equivalent 2SLS estimates. The regression for average wages in columns 1 and 4 are estimated conditional on an individual being employed in a job covered by social security at least one day during the year, daily wages in columns 2, 3, 5, and 6 are measured on June 30 of the relevant year and condition on full- or part-time employment on that day. All coefficients are estimated using the specification defined in Equation (2), wages are deflated to 2010 values. Standard errors are clustered by initial district. +  $p < 0.1$ , \*  $p < 0.05$ , \*\*  $p < 0.01$ .



## 5 Mechanisms and interpretation

Having established that the conational share in the first job an immigrant holds has a negative effect on subsequent employment rates, I now turn to understanding the mechanisms that underpin this result. I close by relating the negative employment effect documented here to previous evidence on the effect of neighbourhood ethnic networks.

### 5.1 Wage and productivity in the first job

The conational share in the first job may be associated with hard-to-observe characteristics of the first job. For example, a higher conational share is correlated with having received the job through a referral (Dustmann et al., 2016), or having a conational manager (Åslund et al., 2014). The immigrant share, i.e. both the conational share and the other migrant share, might also be negatively correlated with firm productivity (Damas de Matos, 2012). The immigrant share may also directly affect worker productivity in the first job, particularly if there are costs to working in mixed teams (Glover et al., 2017; Hjort, 2014; Lazear, 1999a,b).<sup>10</sup>

In all of these cases, the conational share and, perhaps to a lesser extent, the other immigrant share, will have an effect on the wage in the first job. The wage in the first job might in turn affect the probability of switching jobs, either by affecting a worker's starting position on the job ladder (Burdett and Mortensen, 1998), or by altering a worker's threat point when bargaining (Postel-Vinay and Robin, 2002). Furthermore it may affect unemployment duration conditional on becoming unemployed if it affects unemployed workers' reservation wages, or if, for example, past wages are taken as a signal of productivity.

However, the types of relationships described above are arguably better classified as confounders for the effect of the composition of the set of coworkers, rather than mechanisms, since they are not direct outcomes of the conational or other migrant shares. The use of the IV estimation strategy described in Section 3.2 is in part intended to rule out such confounding by correlated job characteristics. Nevertheless, to check that the IV strategy is working as intended in this respect, I estimate the relationship between the initial conational share and other characteristics of the first job. The results are reported in Table 6. In particular, I consider the effect of the conational and other migrant shares on the wage in the first job, the duration of the first job, and an indicator for leaving the first job for another job (as opposed to unemployment). The OLS estimates suggest that the conational share in the first job is positively associated with the wage, but negatively associated with the job duration and negatively associated with the probability of leaving the first job for another form of employment. The IV estimates, on the other hand, while imprecise, do not suggest a significant effect of the conational share on any of these outcomes. A ten-percentage-point increase in the initial conational share raises wages by 1.2 percent, reduces the duration of the first job by 4.3 percent and reduces the probability of an employment to employment transition by 1 percentage point, with all three effects statistically

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<sup>10</sup>Peri and Sparber (2009) and Ottaviano and Peri (2012), provide evidence of aggregate complementarities between immigrants and natives, however it is not clear whether such complementarities arise within firms, or by increasing the scope for specialisation across firms.

insignificant. On the other hand, there is some evidence that the other immigrant share has a statistically significant effect on wages, which increase 3.1 percent, and the probability of a job-to-job transition, which decreases by 1.8 percentage points.

Table 6: Outcomes in first job

	ln(wage <sub>0</sub> )		ln(duration)		EE transition	
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	2SLS	OLS	2SLS	OLS	2SLS
$s_i^{own}$	0.19** (0.037)	0.12 (0.25)	-0.41** (0.13)	-0.42 (0.45)	-0.076** (0.022)	0.096 (0.14)
$s_i^{other}$	0.18** (0.053)	0.31* (0.14)	-0.50** (0.094)	0.45 (0.31)	-0.054** (0.018)	-0.18* (0.085)
Observations	39069	39069	37937	37937	37768	37768
KP F-statistic		17.2		17.0		17.0
Subsample	all	all	UT	UT	UT	UT

*Note:* Results from cross-sectional estimates of the relationship between shares in the first job and other characteristics of the first job. UT denotes untruncated job spells, i.e. completed job spells. Standard errors clustered by initial district. + p<0.1, \* p<0.05, \*\* p<0.01

## 5.2 Job search and social networks

Given that the effect of the initial conational share does not appear to be mediated by other characteristics of the first job, I turn now to the effect of the initial conational share on subsequent job search behaviour. Individuals are known to use their social networks both as a source of information about job openings (Calvó-Armengol and Jackson, 2004; Boucher and Goussé, 2019) and as a source of referrals when applying for jobs (Montgomery, 1991; Galenianos, 2013; Dustmann et al., 2016). Especially relevant in this context, Eliason et al. (2022) show that coworker networks are a particularly important source of referrals, more so than the type of residential networks that have been a focus of previous research on immigrant earnings assimilation (e.g. Edin et al., 2003). Coworkers from an immigrant’s first job are likely to be a particularly important source of information for immigrants’ subsequent job search.

To understand how a higher conational share in an immigrant’s first job affects subsequent job search, I first look at how the conational share affects transitions to new jobs. I regress an indicator for the presence of a coworker of a given type from the first job held in Germany in the firm where an immigrant starts a new job, which I interpret as indicative of the worker receiving a referral, on the share variables of interest. I report the estimates in Table 7. For all transitions into a new job, and for the subset where the immigrant is coming from unemployment, a ten-percentage-point-higher conational share raises the probability of receiving a referral from a past conational coworker by 2 percentage points and lowers the probability of receiving a referral from a past native coworker by one percentage point. The total effect of the conational share on receiving any kind of referral (not shown) is statistically significantly positive.

Workers will endogenously choose to search for a job based on the composition of their networks; the effect of the conational share on the probability of receiving a referral may indicate that immigrants in high-conational share first jobs have a higher-quality network, leading them to use referrals more often.<sup>11</sup> To show that this is not the case, I estimate the same specification on a sample of workers who were displaced from a job by an establishment closure and therefore forced to search for a job (as in, e.g., Cingano and Rosolia, 2012, Glitz, 2017, or Eliason et al., 2022; plant closures are measured using worker outflows, following Hethy-Maier and Schmieder, 2013). For this admittedly small sample, a higher conational share only raises the probability of a referral from a conational by a statistically insignificant 0.6 percentage points and the net effect is zero. These results suggest that a higher conational share in the first job leads immigrants to rely more on past conationals when searching for a job, but that a higher conational share among former coworkers does not imply more productive job finding networks, since the network is no more able to provide jobs when immigrants are unexpectedly forced to search for a job than when the share of conationals in the network is lower.

Table 7: Referrals in subsequent jobs

	P(own ref.)			P(nat. ref.)		
	(1)	(2)	(3)	(4)	(5)	(6)
$s_i^{own}$	0.20** (0.022)	0.21** (0.22)	0.055 (0.11)	-0.091** (0.017)	-0.092** (0.027)	-0.072 (0.11)
$s_i^{other}$	0.046** (0.012)	0.042** (0.014)	0.11+ (0.063)	-0.025+ (0.063)	-0.040** (0.013)	-0.046 (0.014)
Observations	145329	85668	2595	145329	85668	2595
Individuals	25080	21182	1806	25080	21182	1806
$R^2$	0.2	0.2	0.5	0.2	0.2	0.5
Subsample	all	U	C	all	U	C

*Notes:* OLS estimates. Dependent variable is an indicator for presence of a coworker from first job at the start of a subsequent job. U = unemployment to employment transitions; C = unemployment follows plant closure. SE clustered by initial district. +  $p < .1$ , \*  $p < .05$ , \*\*  $p < .01$

The composition of an immigrant’s initial set of coworkers likely affects an immigrant’s subsequent job search by changing the types of ties that make up the job search network. While Granovetter (1995) argued that a higher proportion of weak ties in a network improves search outcomes, Montgomery (1992) noted that this is only true if the job offer rate is higher from weak ties than strong ties and more recent evidence indicates that strong ties are a more productive source of referrals than weak ties (Gee et al., 2017a,b). If we treat conationality as a proxy for being a strong tie, consistent with theories of homophily in social networks (e.g. Mcpherson

<sup>11</sup>Although note that, while immigrants are in general less likely to be employed than natives (OECD, 2020), the evidence in Table A.6 suggested that the conational share is not a straightforward measure of coworker quality, at least as measured by coworkers’ prior attachment to the labour market.

et al., 2001), then the evidence in Table 7 is consistent with the recent evidence on strong ties being more productive at the margin.

While the results in Table 7 show that a higher conational share in the first job leads immigrants to rely more on past coworkers for subsequent jobs, it doesn't tell us how this increased reliance affects the outcomes of immigrants' job search. In Table 8, I show that a higher conational share lowers the probability of moving to another job when a job spell ends. Conditional on becoming unemployed, a higher conational share also appears to weakly increase the duration of the unemployment spell. In Table 9 I report OLS and IV estimates for different samples of the effect of the conational share on the duration of unemployment spells, conditional on becoming unemployed.<sup>12</sup> While the coefficient is only significant for the OLS estimates on the full sample, the point estimate is nevertheless quite stable across the different specifications. The effect of the other immigrant share, on the other hand, is much more unstable and varies substantially across specifications.

Table 8: Employment transitions

	OLS	2SLS
$s_i^{own}$	-0.11** (0.017)	-0.23+ (0.12)
$s_i^{other}$	-0.022 (0.015)	-0.029 (0.063)
Observations	170022	170022
Individuals	31088	31008
KP F-statistic		7.0

*Note:* Outcome is an indicator from moving from a job to another job, rather than unemployment, when completing a job spell. Standard errors clustered by initial district. + p<0.1, \* p<0.05, \*\* p<0.01

To summarise, the increased share of conationals in the first job leads to an increased reliance on coworkers when searching for subsequent jobs, but worse search outcomes. I therefore interpret the estimates reported in Tables 7–9 as providing evidence that a higher conational share in the first job lowers an immigrant's subsequent job search capital. Immigrants appear to learn how to search for subsequent jobs in Germany, without having to rely on their network, at least in part through their interactions with natives in the first job.

### 5.3 Human capital accumulation

An alternative explanation for worse job search outcomes, however, is that the conational share in the first job directly slows down the immigrant's accumulation of productive skills. Acquiring

<sup>12</sup>I do not report IV estimates for the subset of individuals who experience plant closures since this group is too small, affecting power.

Table 9: Log unemployment duration

	OLS				2SLS	
	(1)	(2)	(3)	(4)	(5)	(6)
$s_i^{own}$	0.28** (0.097)	0.10 (0.076)	0.39 (0.30)	0.11 (0.35)	0.38 (0.34)	0.39 (0.30)
$s_i^{other}$	-0.058 (0.052)	-0.047 (0.041)	-0.60* (0.27)	-0.64* (0.25)	0.22 (0.32)	0.27 (0.25)
Observations	141393	119902	4024	3287	141393	119902
Individuals	36302	29557	2955	2354	36302	29557
KP F-statistic					9.8	6.9
subsample	all	UT	C	UTC	all	UT

*Notes:* Dependent variable is the log of unemployment duration. UT = untruncated observations, i.e. completed unemployment spells; C = unemployment follows plant closure. SE clustered by initial district. +  $p < .1$ , \*  $p < .05$ , \*\*  $p < .01$

host country-specific human capital has been shown to account for a substantial portion of the convergence of immigrant wages to native wages over time (Eckstein and Weiss, 2004). Furthermore, Battisti et al. (2022) show that a higher share of conationals in the district of residence lowers the acquisition of host country-specific human capital in the longer run. They argue that this is because a larger share of conational co-residents makes job-finding easier, lowering the benefit of acquiring host country-specific human capital. A higher conational coworker share may likewise slow an immigrant’s acquisition of Germany-specific human capital, making them less productive and making it harder for them to find jobs.

The SIEED does not contain information on human capital formation that would allow me to test this possibility, however, the matched SOEP data on non-return migrants can provide some descriptive evidence. In Table 10, I report the estimates from a linear probability model associating the share variables and an indicator for German proficiency, measured at the time of the survey, controlling for premigration characteristics, such as premigration German proficiency. Both the conational and other migrant shares are negatively associated with human capital in the short-run, with a ten-percentage-point increase in either decreasing the probability of being proficient in German by 4.5–5 percentage points. However, this association does not persist into the long-term and is common to both share variables. It therefore cannot explain the negative effect on job search and employment rates that is specific to the conational share.

On the other hand, the conational share is negatively associated with having completed some form of training or education in Germany, while the other migrant share is not (column 2) and this association is entirely due to training that took place after the start of the first job (column 3). This association could, however, be explained by the fact that individuals with reduced employment rates or who have dropped out of the labour market may have fewer incentives to participate in training, if they don’t expect to find a job. Lower employment rates will also directly lower access to on-the-job training, in the form of apprenticeships, an important component of job training in Germany. As a result, while the evidence presented in Table 10

does not conclusively rule out differential Germany-specific human capital accumulation as a mechanism to explain the negative effect of the initial conational share on job-finding ability, it suggests it can at best only explain part of the worse job search outcomes induced by a higher conational share in the initial job.

Table 10: Human capital accumulation

	(1)	(2)	(3)
	Proficiency	Training in DE	Training   entry
$\mathbf{1}(t \in [0, 2]) \times \text{Conat. share}$	-0.50** (0.16)	-0.017 (0.049)	0.043 (0.042)
$\mathbf{1}(t \in [3, 5]) \times \text{Conat. share}$	-0.28 (0.17)	-0.088 (0.056)	-0.040 (0.052)
$\mathbf{1}(t \geq 6) \times \text{Conat. share}$	-0.10 (0.087)	-0.16** (0.057)	-0.14** (0.054)
$\mathbf{1}(t \in [0, 2]) \times \text{Other mig. share}$	-0.45* (0.18)	0.029 (0.054)	0.083* (0.037)
$\mathbf{1}(t \in [3, 5]) \times \text{Other mig. share}$	-0.23+ (0.13)	-0.088 (0.062)	-0.052 (0.045)
$\mathbf{1}(t \geq 6) \times \text{Other mig. share}$	-0.075 (0.079)	-0.039 (0.077)	-0.025 (0.067)
Observations	1687	10061	10061
Individuals	850	863	863
$R^2$	0.28	0.23	0.26

*Note:* The dependent variable in column 1 is an indicator for reporting being proficient in German at time  $t$ , in column 2 it is an indicator for having completed some form of post-school education in Germany by time  $t$ , in column 3 it is an indicator for having completed some form post-school education in Germany that took place after having entered the labour market by time  $t$ . All specifications include controls for pre-migration characteristics, method of finding first job, other job characteristics, and demographic characteristics. Standard errors clustered by initial district. +  $p < 0.1$ , \*  $p < 0.05$ , \*\*  $p < 0.01$

## 5.4 Interpretation in relation to prior research

Recent evidence on the effect of the conational residential network suggests a dynamic trade-off. Immigrants living in areas with more conationals are better integrated into the labour market in the short-run, but these differences disappear in the long-run (Battisti et al., 2022). In a similar vein, Gagliarducci and Tabellini (2022) find that a greater local density of ethnic social organisations, specifically Italian Catholic churches in the US, increases labour force participation but in lower-quality jobs and occupations.

Battisti et al. (2022) suggest their dynamic effect arises because a higher conational share among neighbours, by increasing employment now, lowers the incentive to acquire host-country specific human capital, which crowds out employment now in return for increasing employment in the future. However, the findings reported here suggest another, potentially complementary

reason for the dynamic tradeoff they document. Individuals living in a location with a higher share of conationals may be able to draw on these conationals to find a job more quickly, however these jobs, potentially obtained through referrals, are likely to be in firms with a higher share of conationals. While a higher conational residential share would therefore speed up entry into the labour market, it will slow down convergence to natives once entry takes place.

The individuals in the SIEED are only observed once they find work. However, I do provide supporting descriptive evidence, drawing on the SOEP, for the mechanism described here. In Figure A.6, I plot the average conational and other migrant share by years until the first job. While the sample is small, a relatively clear pattern nevertheless emerges. Individuals who find work quicker do so in higher conational share firms, for which there may be a future cost, in reduced subsequent employment. The share of immigrants from other countries of origin, on the other hand, does not follow such a clear trend.

## 6 Conclusion

In this paper I have shown that starting one's career in an establishment with a high share of conationals has negative long-term effects on an immigrant's labour market outcomes and particularly their employment rate. This is in contrast to the literature on initial residential conditions for newly arrived immigrants, where a high share of conationals in an immigrant's location of residence, by expanding the size of an individual's network, is generally thought to have positive effects on an immigrant's labour market outcomes. The effect is also specific to an immigrant's conationals; there is no statistically significant penalty for working with immigrants from other countries of origin. Descriptive evidence suggests that the negative effect is not due to a reduced acquisition of Germany-specific human capital. Instead, I show indirect evidence that working with more conationals worsens the quality of an immigrant's social network, making it harder to find jobs in future. Future research could move beyond the first job, to understand what role improvements to coworker networks over time spent in the host country play in longer-term immigrant earnings assimilation.

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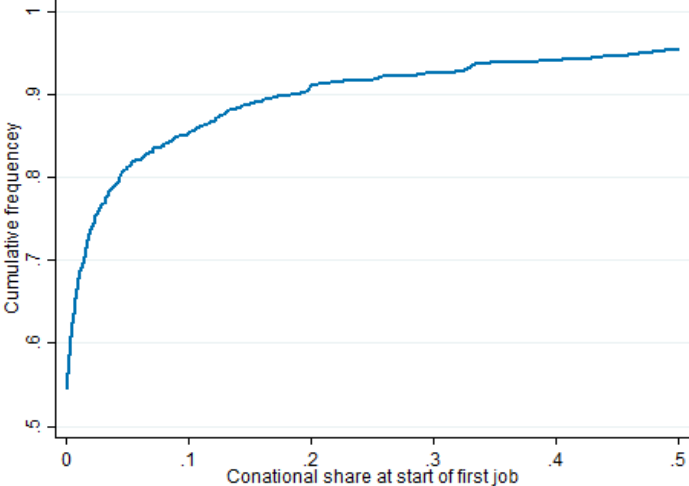
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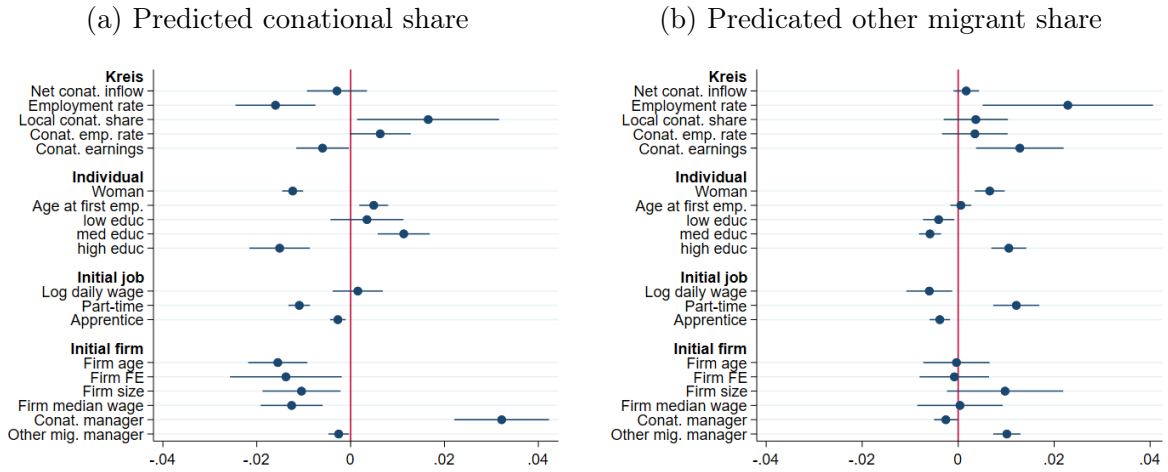
# A Supplementary figures and tables

Figure A.1: CDF of conational share in first job in SOEP



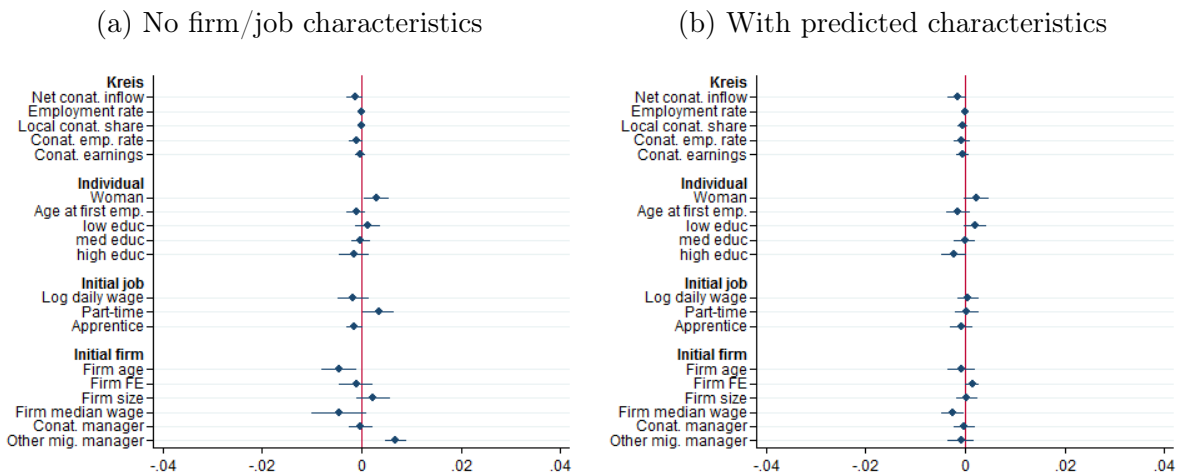
*Notes:* Empirical CDF of the initial conational share in the first job held by an immigrant in my sample. The distribution is truncated at 50, for ease of representation.

Figure A.2: Bivariate correlations of instrument



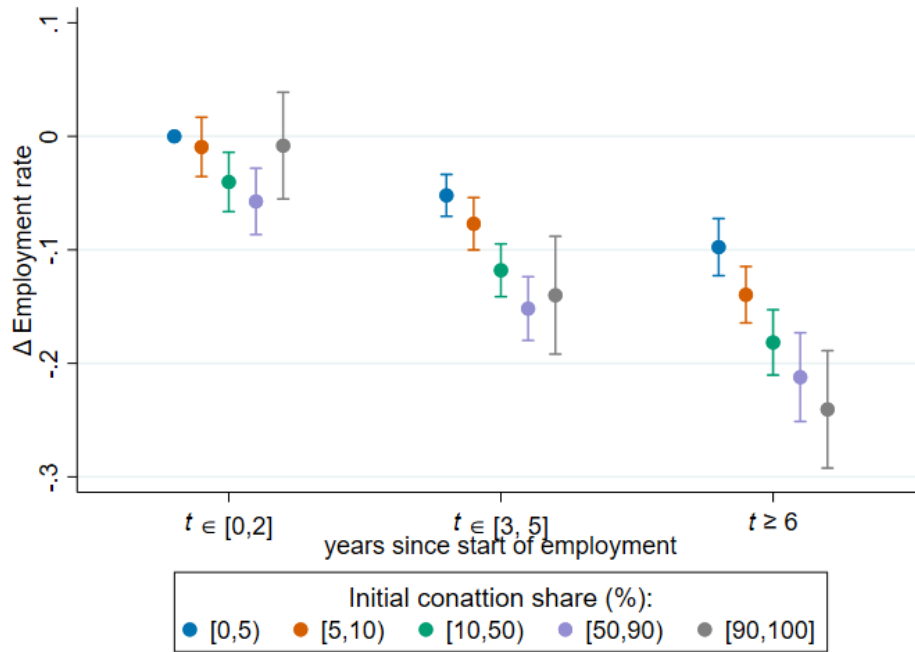
Notes: Bivariate association of instrument with aggregate, individual, and realised firm characteristics. Each association is estimated separately; the dependent variable in each specification has been standardised to have mean 0 and standard deviation 1, while the predicted share is rescaled to lie on [0,100]. Standard errors are clustered by entry district.

Figure A.3: Instrument validity, other migrant share



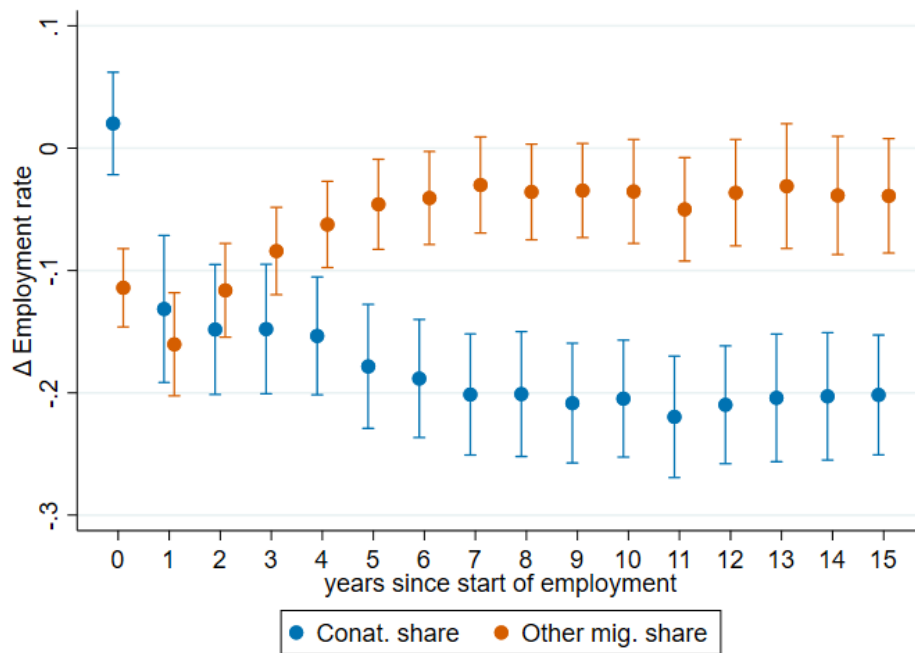
Notes: Effect of predicted share of immigrants from other countries on other characteristics. Each association is estimated separately; the dependent variable in each specification has been standardised to have mean 0 and standard deviation 1, while the predicted share is rescaled to lie on [0,100]. All specifications include labour market  $\times$  nationality  $\times$  entry year and district  $\times$  nationality fixed effects. Standard errors are clustered by entry district.

Figure A.4: Non-linear employment effect of composition of coworkers



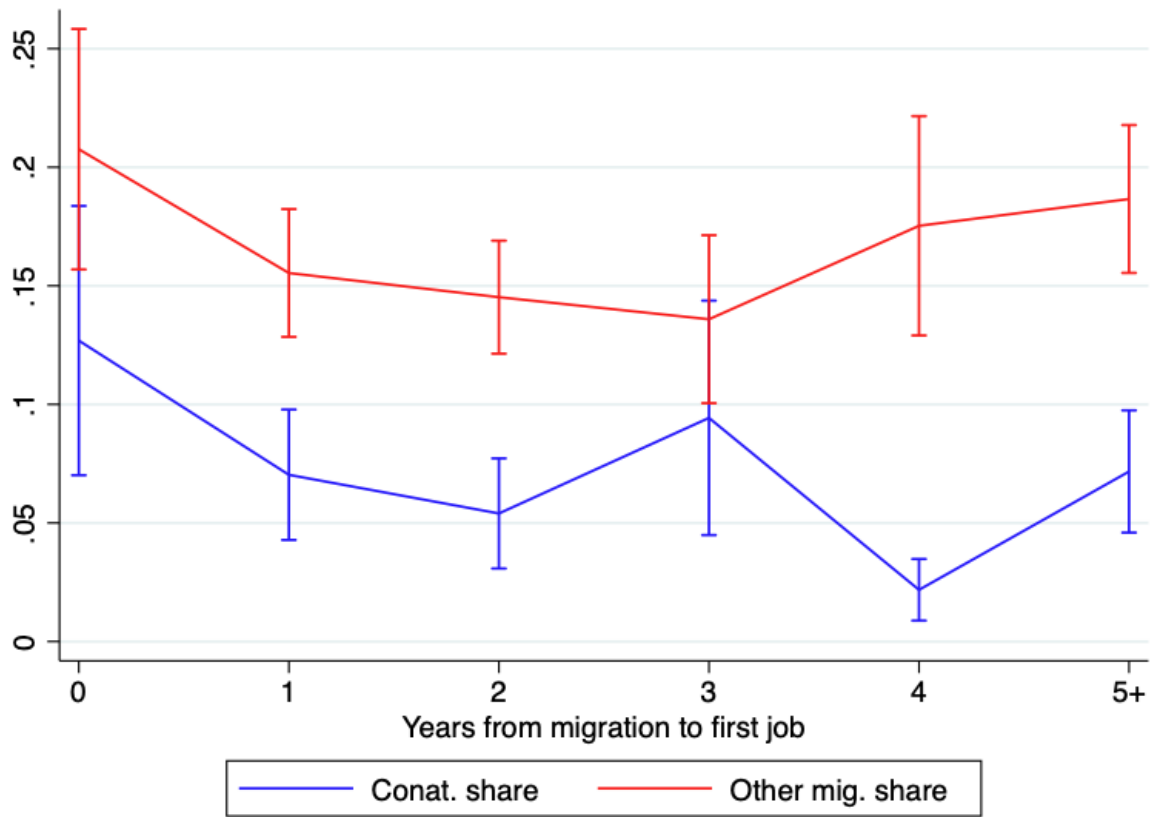
Notes: Indicators for each category, coworker share in  $[0, 5)$  in the first two years of employment is the omitted category. The full set of controls and fixed effects is included, 95 per cent confidence intervals are calculated using standard errors clustered by individual.

Figure A.5: Annual associations



Notes: OLS estimates of the annual association of the conational share other immigrant share. 95 per cent confidence intervals are calculated using standard errors clustered at the initial district level.

Figure A.6: Time taken until first employment and initial share



*Notes:* Mean and 95 per cent confidence intervals for the conational and other migrant share in the first job. N = 863 across all years. Source: IAB-SOEP-MIG-ADIAB.



Table A.1: Country groups, SIIED

	N	Share
Poland	5744	14.70
Yugoslavia, Serbia, Montenegro	4819	12.33
other Asia	2953	7.56
Romania	2324	5.95
Russia, Belarus, USSR	2155	5.52
other Africa	1972	5.05
Croatia	1050	2.69
Portugal	1022	2.62
France	1016	2.60
Hungary	1005	2.57
China	1000	2.56
ex-Czechoslovakia	981	2.51
other America	954	2.44
USA, Canada	946	2.42
Spain	807	2.07
Ukraine, Moldova	802	2.05
Bosnia and Herzegovina	775	1.98
Morocco	775	1.98
Bulgaria	753	1.93
Uk, Ireland	743	1.90
Austria	668	1.71
Iran	627	1.60
Vietnam	624	1.60
India	491	1.26
Netherlands, Luxemburg	468	1.20
Afghanistan	439	1.12
Irak	351	0.90
Albania	334	0.85
Estonia, Latvia, Lithuania	315	0.81
other Europe	268	0.69
Thailand	230	0.59
Macedonia	221	0.57
Ghana	210	0.54
Lebanon	207	0.53
Sri Lanka	206	0.53
Denmark, Sweden	155	0.40
Tunisia	129	0.33
Philippines	126	0.32
Belgium	103	0.26
Ethiopia	91	0.23
Oceania	66	0.17
Slovenia	59	0.15
Switzerland	50	0.13
Finland	35	0.09
Total	39069	100.00

*Note:* Refers to first nationality reported in social security notifications.

Table A.2: Summary statistics, SOEP-IAB data

	Mean	St. dev.	N
Panel A			
Employment rate	0.74	0.38	10061
Annual wage earnings	21256.1	15024.9	7493
$\mathbf{1}(t \in [0, 2])$	0.25	0.44	10061
$\mathbf{1}(t \in [3, 5])$	0.23	0.42	10061
$\mathbf{1}(t \geq 6)$	0.52	0.50	10061
Panel B			
Woman	0.50	0.50	863
Age at migration	29.32	9.04	863
Employed before migrating	0.71	0.46	863
Education	0.14	0.34	863
Low education	0.40	0.49	863
Medium education	0.32	0.47	863
High education	0.29	0.45	863
Support (family)	0.47	0.50	863
Support (friends)	0.10	0.31	863
Support (both)	0.05	0.22	863
No support	0.37	0.48	863
Panel C			
First job through contacts	0.56	0.50	863
Years until first job	3.27	3.02	863
Daily wage	43.1	34.3	863
Firm size	470.4	2221.8	863
Firm median wage	74.3	39.5	863
Firm age	13.0	10.5	863
Conat. share	0.070	0.19	863
Other mig. share	0.17	0.20	863

*Note:* Panel A reports time-varying summary statistics for the years since the first job, average earnings are conditional on being employed on June 30. Panel B reports summary statistics on pre-migration characteristics, including whether an immigrant had any support from someone in Germany when migrating. Panel C reports summary statistics on the characteristics of the first job held after migration and the firm where the job was held. Wages and earnings are deflated and reported in 2010 Euros.

Table A.3: Country groups, SOEP-IAB

	N	Share
Russia	323	37.43
Romania	114	13.21
Poland	93	10.78
ex-Yugoslavia	71	8.23
Turkey	65	7.53
Asia	52	6.03
Italy	41	4.75
Other Europe	38	4.40
Africa	29	3.36
Greece	2*	2.55
Others	//	////
Total	863	100.00

*Note:* Refers to country of birth (as self-reported in the SOEP) for individuals born without German nationality. The table has been censored in accordance with IAB data protection requirements.

Table A.4: First stage effect of predicted other migrant share on the realised share.

	(1)	(2)	(3)	(4)	(5)	(6)
$z_i^{own}$		0.050 (0.040)	0.093* (0.037)	0.063+ (0.037)	0.097+ (0.052)	0.13* (0.054)
$z_i^{other}$	0.69** (0.029)	0.70** (0.030)	0.52** (0.029)	0.53** (0.033)	0.59** (0.045)	0.56** (0.030)
Pred. charact.	No	No	No	No	No	Yes
N	39069	39069	39069	39069	37302	39069
$R^2$	0.16	0.16	0.44	0.52	0.58	0.52
FE	-	-	LxNxY, D	LxNxY, DxN	LxNxY, DxN, DxY	LxNxY, DxN

*Note:* Static first-stage relationship between predicted share of immigrants from other countries,  $z_i^{other}$ , and the realised share of immigrants from other countries in the first job. Included predicted characteristics are part-time status, firm age, conational manager, other migrant manager, log predicted firm size and log predicted median wage. L = labour market, N = nationality, D = district, Y = year of first job. Standard errors clustered by district +  $p < 0.1$ , \*  $p < 0.05$ , \*\*  $p < 0.01$

Table A.5: Effect heterogeneity by individual and firm characteristics

	(1)	(2)	(3)	(4)
$s_i^{own}$	-0.33*			
	(0.13)			
male $\times s_i^{own}$		-0.33*		
		(0.13)		
female $\times s_i^{own}$		-0.34*		
		(0.14)		
low educ $\times s_i^{own}$			-0.34*	
			(0.13)	
med educ $\times s_i^{own}$			-0.36**	
			(0.13)	
high educ $\times s_i^{own}$			-0.20	
			(0.14)	
small $\times s_i^{own}$				-0.46+
				(0.28)
large $\times s_i^{own}$				-0.33*
				(0.13)
$s_i^{other}$	-0.034			
	(0.097)			
male $\times s_i^{other}$		0.0021		
		(0.11)		
female $\times s_i^{other}$		-0.072		
		(0.096)		
low educ $\times s_i^{other}$			-0.060	
			(0.097)	
med educ $\times s_i^{other}$			0.034	
			(0.11)	
high educ $\times s_i^{other}$			0.014	
			(0.19)	
small $\times s_i^{other}$				-0.056
				(0.096)
large $\times s_i^{other}$				0.0048
				(0.12)
Observations	32123	32123	32123	32123
Individuals	32123	32123	32123	32123
KP F	14.2	11.1	9.3	11.6

Notes: Cross-sectional IV estimates, see main text for details. Standard errors are clustered by district. +  $p < .1$ , \*  $p < .05$ , \*\*  $p < .01$

Table A.6: Effect heterogeneity by local and conational characteristics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$s_i^{own}$	-0.33*	-0.42*	-0.35 <sup>+</sup>	-0.30*	-0.43*	-0.25	-0.35*
	(0.13)	(0.18)	(0.19)	(0.14)	(0.19)	(0.19)	(0.17)
Kreis emp. $\times s_i^{own}$					-0.054		
					(0.089)		
Kreis conat. share $\times s_i^{own}$						-0.078	
						(0.063)	
Coworker emp. $\times s_i^{own}$							-0.059
							(0.048)
$s_i^{other}$	-0.034	-0.020	0.051	-0.027	-0.020	0.027	-0.032
	(0.097)	(0.12)	(0.12)	(0.097)	(0.12)	(0.12)	(0.10)
Kreis emp. $\times s_i^{other}$					-0.044		
					(0.084)		
Kreis conat. share $\times s_i^{other}$						0.13	
						(0.096)	
Coworker emp. $\times s_i^{other}$							-0.056
							(0.072)
Kreis emp.		0.0087			0.026		
		(0.021)			(0.036)		
Kreis conat. share			0.0074			-0.011	
			(0.016)			(0.031)	
Coworker emp.				0.017			0.035*
				(0.020)			(0.015)
Observations	32123	23933	19529	32123	23933	19529	32123
KP F-statistic	14.2	9.0	6.5	7.7	7.1	5.6	5.6

*Notes:* Cross-sectional IV estimates. All estimation-specific controls and interaction variables have been standardised to have mean zero and standard deviation one. Columns 2 and 5 include Kreis-level employment rates as a control, calculated using data from the *Mikrozensus* and regional statistical offices, available from 1995. Columns 3 and 6 include the conational share in the Kreis as a control, calculated using data from the *Ausländerzentralregister* and the *Mikrozensus*, available from 1998. Columns 4 and 7 include controls for the employment rate of coworkers in the five years preceding the first job, calculated from the SIEED. All specifications include labour market by nationality by starting year and district by nationality fixed effects. Standard errors are clustered by district. +  $p < .1$ , \*  $p < .05$ , \*\*  $p < .01$

## B The sign of the bias induced by selective return migration

The IAB-SOEP Migrant Sample is made up of survivors, immigrants who were still in Germany in 2013 and 2014 in order to be interviewed. It is generally accepted that return migrants had worse labour market outcomes, summarised by earnings, before returning than immigrants who stay (Borjas, 1985; Lubotsky, 2007; Sarvimäki, 2011). This tells us that earnings have a negative effect on return migration, or that return migration and earnings share some common unobservable cause—return migrants might be intrinsically less productive individuals—either of which can bias estimates of the rate of earnings convergence of immigrants to natives over time (Abramitzky et al., 2014). However, when studying the effect of an initial condition, the conational share in the first job, on subsequent labour market outcomes, the sign of the selection bias will depend not only on the effect of earnings on return migration, but also on any effect the initial conational share might have on return migration.

To focus on intuition and to emphasise the fact that the bias induced by selective return migration is independent of the bias induced by selection into treatment on unobservables, I derive the sign of the selection bias under the simplifying assumption that (i) the initial conational share,  $S$  is randomly assigned; and (ii) there are no systematic determinants of subsequent employment rates  $Y$  besides  $S$ . Furthermore, assume that the conational share is either low or high, i.e.  $S \in \{0, 1\}$ . Assuming the effect of  $S$  on  $Y$  is linear, the structural equation for  $Y$  is simply:

$$Y = a + \beta S + \varepsilon_Y. \quad (4)$$

The structural error term  $\varepsilon_Y$  is mean-zero<sup>13</sup> and independent of  $S$ , since there is no confounding. To model selection, I assume that latent utility  $C^*$  is a linear function of  $S$ ,  $Y$ , and a mean-zero structural error term:

$$C^* = \alpha_S S + \alpha_Y Y + \varepsilon_{C^*}, \quad (5)$$

where  $\alpha_i \in \mathbb{R}$ ,  $i \in \{Y, S\}$ . An individual is assumed to return migrate,  $C = 1$ , if latent utility is below some fixed threshold:

$$C(S, Y) = \begin{cases} 1 & \text{if } C^* < K, \\ 0 & \text{otherwise.} \end{cases} \quad (6)$$

Equation (6) captures the fact that  $C$  is endogenously determined by both  $S$  and  $Y$ . The sign of  $\alpha_i$ ,  $i \in \{Y, S\}$ , encodes hypothetically testable assumptions about the effect of the observable variables  $Y$  and  $S$  on  $C$ . I now show how the selection bias from conditioning the analysis on  $C = 0$  depends on the signs of  $\alpha_S$ ,  $\alpha_Y$ , and  $\beta$ . Since the structural equation is linear and  $S$  is assumed to be randomly assigned, the true parameter of interest,  $\beta$ , can be defined as

$$\beta = \frac{\text{Cov}(Y, S)}{\text{Var}(S)} \quad (7)$$

---

<sup>13</sup>Furthermore, we must have  $\varepsilon_Y \in [-a, 1 - (a + \beta)]$ , since  $Y \in [0, 1]$

Since we only observe individuals with  $C = 0$ , however, the OLS estimand on this restricted sample is

$$\begin{aligned}
\hat{\beta} &= \frac{\text{Cov}(S, Y|C = 0)}{\text{Var}(S|C = 0)} \\
&= \beta + \frac{\text{Cov}(S, \varepsilon_Y|C = 0)}{\text{Var}(S|C = 0)} \\
&= \beta + \frac{\text{Cov}(S, \varepsilon_Y|C^* \geq K)}{\text{Var}(S|C^* \geq K)}
\end{aligned} \tag{8}$$

The sign of the bias induced by conditioning on the endogenous variable  $C$  will therefore depend on the sign of the conditional covariance of  $S$  and  $\varepsilon_Y$ , since the conditional variance of  $S$  is positive. Note that  $\text{Cov}(S, \varepsilon_Y) = 0$  in the full sample by assumption, but not in the restricted sample of non-return migrants. The sign of the conditional covariance can be calculated as

$$\begin{aligned}
&\text{Cov}(S, \varepsilon_Y|C^* \geq K) \\
&= \text{E}[S\varepsilon_Y|C^* \geq K] - \text{E}[S|C^* \geq K]\text{E}[\varepsilon_Y|C^* \geq K] \\
&= \text{E}[\varepsilon_Y|C^* \geq K, S = 1]\Pr(S = 1|C^* \geq K)
\end{aligned} \tag{9}$$

$$\begin{aligned}
&\quad - \text{E}[S|C^* \geq K]\text{E}[\varepsilon_Y|C^* \geq K] \\
&= \{\text{E}[\varepsilon_Y|C^* \geq K, S = 1] - \text{E}[\varepsilon_Y|C^* \geq K]\}\Pr(S = 1|C^* \geq K),
\end{aligned} \tag{10}$$

where the second equality follows from the law of iterated expectations and the third from the fact that  $S$  is a Bernoulli random variable, so its expectation is the probability that  $S = 1$ . The sign of the conditional covariance will depend on the sign of the difference of the two conditional expectations in parentheses in Equation (10),  $\text{E}[\varepsilon_Y|\cdot]$ . Note, however, that  $\varepsilon_Y$  is a mean-zero random variable and that its distribution is truncated when calculating the expectations  $\text{E}[\varepsilon_Y|\cdot]$ . The sign of the conditional expectations will therefore depend on whether the right or the left tail of the distribution is truncated. Furthermore, the difference between the expectations will depend on which distribution is more severely truncated. The truncation condition  $C^* \geq K$  can be re-written

$$\alpha_Y \varepsilon_Y \geq K - (\alpha_S + \alpha_Y \beta)S - \alpha_Y a - \varepsilon_{C^*}, \tag{11}$$

This inequality makes clear how the sign of the bias of  $\hat{\beta}$  with respect to  $\beta$  will depend not only on (i) the total effect of employment on return migration, captured by  $\alpha_Y$ ; but also potentially on (ii) the total effect of the conational share on return migration, that is without netting out the part of the effect that is mediated by employment, i.e.  $\alpha_S + \alpha_Y \beta$ . Intuitively, the sign of  $\alpha_Y$  determines whether the distribution of  $\varepsilon_Y$  is left- or right-truncated, and the sign of  $\alpha_S + \alpha_Y \beta$  determines whether the distribution is more or less severely truncated when  $S = 1$ . If both  $\alpha_Y$  and  $\alpha_S + \alpha_Y \beta$  are of the same sign, the bias will be negative, while if  $\alpha_Y$  and  $\alpha_S + \alpha_Y \beta$  are of opposite signs, the bias will be positive.

To see this, note that if  $\alpha_Y > 0$ , the condition  $C^* \geq K$  truncates the left tail of the distribution of  $\varepsilon_Y$ ; the expectations in Equation (10) will be positive. Furthermore, if  $\alpha_S + \alpha_Y \beta >$

0, then the supplementary condition  $S = 1$  truncates the distribution less severely than when the condition is not imposed, since  $S \in \{0, 1\}$ . As a result, we will have

$$E[\varepsilon_Y | C^* \geq K, S = 1] < E[\varepsilon_Y | C^* \geq K] \quad (12)$$

and the bias will be negative. If, on the other hand,  $\alpha_Y < 0$ , the right tail of the distribution is truncated and the expectations in Equation (10) are negative. If  $\alpha_S + \alpha_Y \beta > 0$ , the supplementary condition  $S = 1$  again means the distribution is less severely truncated, implying now that

$$E[\varepsilon_Y | C^* \geq K, S = 1] > E[\varepsilon_Y | C^* \geq K] \quad (13)$$

and the bias will be positive.

An interesting special case arises when the true effect of interest  $\beta = 0$ . Now the gross effect of the conational share on return migration is simply the direct effect,  $\alpha_S$ . In this case, if  $\alpha_Y$  and  $\alpha_S$  are of the same sign, then  $\hat{\beta} < 0$ , while if they are of opposite signs, then  $\hat{\beta} > 0$ . Therefore, if the estimated  $\hat{\beta} < 0$  and one has reason to believe that  $\alpha_Y$  and  $\alpha_S$  are of opposite signs, then the observed association cannot be entirely explained by selection into return migration; it must be that  $\beta < 0$ .

The estimates on dropping out of the sample, using the SIEED, reported in Table 4, suggest that a higher conational share increases return migration, i.e.  $\alpha_S < 0$ . Assuming that any selection bias is not so great as to flip the sign of the employment effect, then  $\beta < 0$ . Furthermore, evidence on the effect of employment on return migration suggests  $\alpha_Y > 0$  (Sarvimäki, 2011), implying that  $\alpha_S + \alpha_Y \beta < 0$ . Selection bias would imply that  $\hat{\beta} < \beta$  in the IAB-SOEP Migrant Sample.