Labor Supply Shocks, Native Wages, and the Adjustment of Local Employment*

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Abstract

By exploiting a rare commuting policy that led to a sharp, sudden, and unexpected inflow of Czech workers to areas along the German-Czech border, we examine the impact of an exogenous immigration-induced labor supply shock on local wages and employment of natives in the policy's immediate aftermath. To guide our empirical analysis, we develop a simple model that – other than the existing literature – allows for heterogeneity in labor supply elasticities or wage rigidities that different groups of workers face. Our results show that the labor supply shock leads to a moderate decline in local native wages and a sharp decline in local native employment. This employment response is almost entirely driven by diminished *inflows* of native workers into work rather than *outflows* into other areas or non-employment, suggesting that "outsiders" shield "insiders" from the increased competition. We also document large *wage* responses for young but strong *employment* responses for older natives, which we attribute to either a larger labor supply elasticity or a higher degree of wage rigidity for the latter. Such responses, while in line with our model, cannot be reconciled with a standard model where labor supply responses are homogeneous across workers.

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I. Introduction

Although numerous papers address the effect of immigration on the wages and employment of native workers, there is still little consensus on whether an immigration-induced labor supply shock has adverse impacts or on the most suitable methodology for addressing this issue.[1] Building on this literature, we revisit this question by exploiting a policy which has been implemented 14 months after the fall of the Berlin wall and which allowed Czech workers to seek employment in eligible German border municipalities but denied residence rights, thereby inducing daily commuting across the border. This commuting policy resulted in an almost ideal exogenous labor supply shock that was unexpected, sudden, and of considerable magnitude, averaging to about 10% of local employment in municipalities closest to the border. The commuting requirement created exogenous variation in the impact intensity at a disaggregated geographic (i.e., municipal) level, which distinguishes our work from other studies that use an experimental design.[2]

A further distinguishing feature of our work is the exceptionally high quality data we have available, which is of longitudinal nature and covers the entire workforce. This allows analysis not only of the short-term effects of native responses for detailed groups of workers (e.g., young unskilled natives) but also of different *types* of employment adjustments. For example, although native employment adjustments in response to an immigration-induced supply shock are typically interpreted as outflows into non-employment, they could also result from fewer non-employed

[1] See, for example, Grossman (1982), Altonji and Card (1991), Goldin (1994), Borjas, Freeman, and Katz (1996, 1997), Card (2001), Borjas (2003), Angrist and Kugler (2003), Manacorda, Manning, and Wadsworth (2012), and Ottaviano and Peri (2012).

[2] See, for example, Card (1990), Hunt (1992), Carrington and Lima (1996), Friedberg (2001), Glitz (2012), and Prantl and Spitz-Oener (2014).

workers entering employment in the affected area.[3] Similarly, adjustments in local employment need not only stem from movements into and out of non-employment; they could also be due to geographic movements across local labor markets, a mechanism found to be essential to explain the long-run effects of adverse demand shocks in the U.S. (see Blanchard and Katz, 1992).[4] To throw more light on these aspects, we provide evidence on the magnitude of each type of response and show how their relative importance varies across worker groups. Thus, the combination of a highly informative policy, a clean identification strategy, and high quality longitudinal data on potentially affected workers allows us to produce a more complete picture of the effects of labor supply shocks than what so far reported.

To structure our empirical analysis and better understand our estimates, we begin with a simple production model of the type standard in the migration literature. We first extend that model by allowing for heterogeneity in natives' employment responses that is either due to different groups exhibiting different local labor supply elasticities, or due to partial wage rigidity in the short run that varies across demographic groups. We show that such heterogeneity may lead to "perverse" wage effects in the sense that the skill group that experiences the largest labor supply shock is not necessarily the skill group whose wages decline the most.[5] Nevertheless, precisely because of the high labor supply elasticity, or the high degree of wage rigidity, this group is likely to exhibit a strong employment response so that relative wage and employment effects between

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[3] An exception is Cohen-Goldner and Paserman (2006) who distinguish, like us, between the effect of immigration to Israel on inflows and outflows from employment using the rotating panel feature of the Israeli labor force survey.

[4] Blanchard and Katz (1992) find that U.S. states that experience an adverse demand shock never fully recover in terms of employment, but that unemployment and wages adjust because of workers moving out of affected states. In the migration literature, however, the question of whether and to what extent an immigration-induced labor supply shock may lead some of the existing workforce to relocate remains controversial (see, e.g., Borjas, Freeman, and Katz, 1997; Filer, 1992; Card, 2001; Butcher and Card, 1991; Card and DiNardo, 2000; and Borjas, 2003, 2006).

[5] Piyapromdee (2014) makes a related point by suggesting that a mainly unskilled immigration shock to a particular area may be exacerbated in its effect on unskilled natives if these are relatively immobile across areas in comparison to skilled natives who are complementary in production.

skill groups need not move in the same direction. This observation underscores the need to analyze immigration-induced wage and employment responses *jointly,* as isolated estimates of wage or employment effects may not only misrepresent the overall impact of immigration (as emphasized by Borjas, 1999), but also its distributionary effects.[6]

Our empirical estimates show that the inflow of Czech workers leads to a moderate decline in local wages and a sharp decline in local employment of natives. Three years into the policy, a 1 percentage point increase in the overall employment share of Czech workers had decreased local native wages by about 0.13 and local native employment by about 0.9 percent.[7] Both responses were remarkably rapid, with the wage response preceding the full employment response. In light of the strong employment response, it is not surprising that the public reaction to the commuting policy became less favorable, which eventually led to a tightening of the policy.[8]

As it is the case for any immigration episode, our findings have to be interpreted in light of the particular policy considered. There are several reasons for why the inflow of immigrants may have led to more adverse effects on natives in ours than in other situations. First, unlike in many other contexts, commuting workers did not live and consume in the affected areas, thus reducing possible demand effects induced by immigrant consumption.[9] Second, it focuses on the

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[6] Most papers in this line of research focus on wage responses only, although Card (1990, 2001, 2007), Altonji and Card (1991), Dustmann, Fabbri, and Preston (2005), Boustan, Fishback, and Kantor (2010), Wagner (2010), and Glitz (2012) consider wage and employment responses. These papers do not, however, investigate how wage and employment responses interact with each other.

[7] Glitz (2012) and Aydemir and Kirdar (2014), using quasi-natural experiments, also find large employment effects, although their specification is not directly comparable to ours. Using a not dissimilar design to ours, Doran, Gelber, and Isen (2015) conclude that the causal impact of extra H-1B visas crowds out employment of other workers in the receiving firm.

[8] See for example *Süddeutsche Zeitung*, 12.6.1992 ("*Immer mehr Arbeitskräfte kommen aus Osteuropa*") and 21.6.1994 ("*Einpendler belasten den Arbeitsmarkt*").

[9] Despite studying cases when immigrants live and consume in the affected areas, most empirical papers address only the production side and do not investigate the impact of immigrant consumption on native-born wages, although some discuss this possibility. For instance, in an early paper, Greenwood and Hunt (1984) suggest that immigration can increase aggregate demand, while Altonji and Card (1991) and Borjas (2013) consider immigrant consumption in their model but not in their empirical analysis. Hercowitz and Yashiv (2002) and Bodvarsson, van den Berg, and Lewer

short-term effects of an unexpected and exceptionally large labor supply shock, affecting a region that had not experienced large immigrant inflows or labor supply shocks in the recent past.[10] Third, the labor supply shock may possibly have been viewed as temporary by firms, making them reluctant to expand capital in response to the shock.

Our decomposition of the overall native employment response into different types of adjustment sheds new insight to the interpretation of employment responses to immigration. First, native employment decreases predominantly through reductions in *inflows* into local employment, whereas *outflows* from the incumbent native workforce are much smaller. This observation indicates that "outsiders" (i.e., workers not employed in the affected area) bear most of the burden of the labor supply shock and thus shield "insiders" (i.e., workers employed in the affected area) from the adverse effects of the shock—either because "outsiders" are particularly elastic in their employment response or because "insiders" are, at least in the short run, protected by partial wage rigidity and firing restrictions. Second, even in the short run, roughly one third of the *local* employment response results from geographic movement to and from employment in other areas not affected by the labor supply shock, meaning that it does not necessarily reflect a reduction in the *overall* employment level.

In terms of differential effects by skill, the inflow of Czech workers leads to larger wage and employment declines for unskilled than skilled natives, which, given Czech workers' lower

(2008) use model-based approaches to reexamine mass migration to Israel and the Miami boatlift, respectively, and conclude that demand effects may delay or abate wage and employment effects on natives.

[10] This distinguishes our border region from e.g. the Miami labor market analyzed in Card (1990), which had a long history of immigration (with 35.5% foreign born). Card (1990) points out that as a result, the "industry distribution in Miami in the late 1970s was well suited to handle an influx of unskilled immigrants", with "textile and apparel industries particularly prominent" (p. 256). Similar considerations hold for studies that exploit historical spatial variation in immigrant inflows for identification, motivated by Bartel's (1989) observation that immigrants tend to settle in areas where others from their country have settled in earlier years. Although it may be plausible to assume that past settlement patterns are otherwise unrelated to changes in future wages and employment, areas with historically large immigration inflows might be quicker to adjust capital and thus more adaptive to labor supply shocks.

level of skills relative to German workers, is in line with the standard immigration model. Breaking wage and employment responses out by age group, however, produces evidence consistent with "perverse" effects. That is, although most Czechs who enter the West German border areas are middle-aged, within skill groups, it is workers under 30 who suffer the largest *wage* decline and those over 50 who suffer the largest *employment* decline. This pattern is inconsistent with standard models of immigration but can be accounted for by a model that allows for a larger employment response (either due to a larger local labor supply elasticity, or a higher degree of wage rigidity) for older than for young workers.

Interestingly, the overall patterns of adjustment to the immigration-induced labor supply shock documented in this paper closely mirror the labor market adjustments in a recession: The business cycle literature highlights that in a recession wages in ongoing jobs are relatively sticky whereas employment drops sharply, which—just like in our case—is mostly accounted for by reduced hiring and not by increased separations (e.g., Hall 2005, Shimer 2005, Rogerson and Shimer 2011, Shimer 2012). This suggests our findings have implications beyond the immigration literature and generally help us to better understand how labor markets respond to shocks.

II. An Equilibrium Model with Heterogeneous Labor Supply and Wage Rigidities

To aid the interpretation of our empirical findings, we commence by setting out a simple model that links immigration-induced labor supply shifts to the employment and wage responses of natives in the local labor market. We assume that (as it is the case in our empirical application) the local labor market under consideration is small relative to the national labor market. In consequence, the change in equilibrium wages (and native employment) in other areas will be negligible even if natives respond to the labor supply shock by moving away from affected areas. We start out with a fully competitive labor market as a benchmark (Section II.B.i), and allow for wage rigidities in a second step (Section II.B.ii). One important distinguishing feature of our model relative to other models is that we allow the labor supply responses of natives, or the degree of wage rigidity, to vary across skill or other demographic groups. Unlike the work by for example Manacorda, Manning and Wadsworth (2012) or Ottaviano and Peri (2012), the model is merely aiding the interpretation of our parameter estimates, and we do not attempt to estimate its structural parameters.

II.A. Basic Set-up

II.A.i. Production

Supposing that output Q in a specific area is produced by combining labor L and capital K according to a Cobb-Douglas production function, then

$$Q = AK^{\alpha}L^{1-\alpha}.$$

Here, labor L is a CES aggregate of unskilled (U) and skilled (S) labor L_g, $g = U, S$:

$$L = \left[\theta_U L_U^{\beta} + \theta_S L_S^{\beta}\right]^{\frac{1}{\beta}},$$

where $\theta_U + \theta_S = 1$, and the elasticity of substitution between the two skill groups equals $\sigma = \frac{1}{1-\beta}$, with $\beta \leq 1$.

Within each skill group g, natives (or incumbents, denoted by L_g^N) and immigrants (or entrants, denoted by L_g^I) are perfect substitutes in production, so that $L_g = L_g^N + L_g^I$.[11] Without loss of

[11] We investigate below wage- and employment responses for different skill groups to the *overall* labor supply shock induced by the commuting policy. This means that in our estimation procedure, we do not allocate Czech workers to skill groups based on their observed skills. Whether Czechs compete (and therefore are substitutes for) natives in a particular skill group will be part of the parameter that we estimate.

generality, we further assume that (as in our empirical setting) there are no immigrants in the base period.

II.A.ii. Labor demand

Assuming that firms are price takers in the labor, capital and product market and normalizing the price of the output good to 1, firms choose labor and capital such that marginal costs equal the marginal products of labor and capital:

$$log\, w_g = log[(1-\alpha)A] + \alpha[log\, K - log\, L] + log\, \theta_g + (\beta - 1)\left[log\, L_g - log\, L\right] \quad \text{(1a)}$$

$$log\, r = log\, \alpha A + (\alpha - 1)[log\, K - log\, L]. \quad \text{(1b)}$$

Suppose that the local supply of capital depends on the rental price of capital in the local labor market under consideration (*r*) and on rental prices in other local markets (***r'***), $K = h(r, \boldsymbol{r'})$, and let λ denote the inverse of the local elasticity of capital with respect to its price r (i.e., $\frac{1}{\lambda} = \frac{\partial h}{\partial r}\frac{r}{h}$).

In Online Appendix A.1, and following Dustmann, Frattini and Preston (2013), we derive the firm's change in the demand of native workers (net of immigrant workers) from skill group g, $dlogL_g^N$, to a *total* immigration-induced labor supply shock (as opposed to the *skill-specific* shock typically considered in the literature) relative to native equilibrium employment in the base period (in head counts), $dI = \frac{dL^I}{L^N}$, resulting in

$$\frac{dlogL_g^N}{dI} = \frac{\varphi s_g + (\beta-1)s_{g'}}{(\beta-1)\varphi}\frac{\mathrm{dlog}w_g}{dI} - \frac{\left(\varphi-(\beta-1)\right)s_{g'}}{(\beta-1)\varphi}\frac{dlogw_{g'}}{dI} - \frac{\pi_g^I}{\pi_g^N} \quad (2)$$

where g' denotes the *other* skill group, $\varphi = -\frac{\alpha\lambda}{1-\alpha+\lambda}$ is the slope of the aggregate labor demand curve, π_g^N and π_g^I denote the share of workers of skill group g (in head counts) among immigrants and natives (i.e., $\pi_g^N = \frac{L_g^N}{L_U^N+L_S^N}$ and $\pi_g^I = \frac{L_g^I}{L_U^I+L_S^I}$), and s_g denotes the contribution of labor type g to the total labor aggregate (see Online Appendix A.I for details).

Suppose that g indexes unskilled labor and g' skilled labor and that immigration is predominantly unskilled (i.e. $\frac{\pi_g^I}{\pi_g^N} > 1$). Equation (2) first illustrates that in the absence of any wage response to immigration (i.e., $\frac{\mathrm{dlog}w_g}{dI} = \frac{dlogw_{g'}}{dI} = 0$), unskilled native employment declines by the rate $\frac{\pi_U^I}{\pi_U^N}$, the relative density of immigrants to natives among unskilled workers. Equation (2) further highlights that a decline in the wage of unskilled labor in response to immigration (i.e., $\frac{\mathrm{dlogw_U}}{dI} < 0$) will dampen the employment response of the unskilled, as both the slope of the demand curve φ and $\beta - 1$ are negative (i.e., $\frac{\varphi s_g + (\beta-1)s_{g'}}{(\beta-1)\varphi} < 0$). Further, the impact of the overall immigration shock on skilled wages is ambiguous (i.e., $\frac{dlogw_{g'}}{dI} \lesseqgtr 0$).[12] Similarly, the impact of an increase in skilled wages on the demand for unskilled native labor is also ambiguous (i.e., $\frac{(\varphi-(\beta-1))s_{g'}}{(\beta-1)\varphi} \lesseqgtr 0$), depending on the response of capital and the degree of substitutability between the different input factors.

II.B. Equilibrium

II.B.i. Competitive Equilibrium with Fully Flexible Wage

In a competitive equilibrium, quantities supplied must equal quantities demanded, and the intersection of the demand curve given by Equation (2) and the supply curve determine the skill-specific and aggregate wages and employment in the local labor market. Using N_g to denote the

[12] There are two opposing forces: skilled wages decrease because of imperfect elasticity of capital, but increase because of imperfect substitution between skilled and unskilled workers.

(fixed) number of natives who could potentially supply labor to the local labor market, the local labor supply function for skill group g is

$$L_g = L_g^I + L_g^N = L_g^I + N_g\ f_g(w_g, \boldsymbol{w'_g}), \tag{3}$$

where immigrants (i.e., new entrants) are (as in Borjas 2013) assumed to supply their labor inelastically, but the local labor supply of natives (i.e., incumbents) depends on skill-specific wages in the market under consideration (w_g) and other local labor markets ($\boldsymbol{w'_g}$). The local labor market elasticity for natives, which we allow to vary by skill group, is then given by $\eta_g = \frac{\partial (N_g\, f_g)}{\partial w_g}\frac{w_g}{N_g\, f_g}$. We must emphasize that this elasticity differs from the elasticities typically estimated in the labor supply literature, which measure the response of individuals to changes in net wages affecting the *national* labor market.[13] We, in contrast, consider a manipulation of *local* labor market conditions to which natives may respond not only by moving into and out of non-employment but also by moving away from, or no longer moving into, the area.[14] The local labor supply elasticity therefore summarizes various potential adjustment mechanisms, such as the internal migration of workers between areas, or entries into and exits from the labor force. These adjustment margins may have different importance for different types of workers and thus help explain why some groups respond more elastically than others. For instance, the local labor supply of skilled workers may be more elastic than that of unskilled workers since the former exhibit higher geographic mobility rates than the latter.

From the labor supply function (3), it follows that

[13] See, e.g., MaCurdy (1981) and Chetty et al. (2011) who estimate the labor supply elasticity at the intensive margin or Blundell, Bozio, and Laroque (2011) who estimate the elasticity at the extensive margin.
[14] Heterogeneity in geographical mobility may have different reasons. For instance, Notowidigdo (2013) shows that labor demand shocks may lead to differential mobility responses for low and high skilled workers because they lead to changes in house prices and transfer payments.

$$dlogw_g = 1/\eta_g dlogL_g^N. \tag{4}$$

By substituting this expression (for both skill groups) into Equation (2) and rearranging, we derive the equilibrium employment response as (see Online Appendix A.II for details):

$$dlogL_g^N = \frac{\eta_g(\beta-1)\left[\frac{\pi_g^I}{\pi_g^N}\left(1-\varphi\eta_{g'}\right)-\Pi(1-\frac{\varphi}{\beta-1})\right]}{1-(\beta-1)\left[\eta_g\left(1+s_g\phi\right)+\eta_{g'}\left(1+s_{g'}\phi\right)-\eta_g\eta_{g'}\varphi\right]}dI, \tag{5}$$

because $\beta \leq 1$, the denominator in (5) will always be positive. The numerator is the difference between the relative density of immigrants to natives in skill group g, $\frac{\pi_g^I}{\pi_g^N}$ and the (weighted) average of these densities in the different skill groups, $\Pi = s_U \frac{\pi_U^I}{\pi_U^N} + s_S \frac{\pi_S^I}{\pi_S^N}$, both weighted by expressions that depend on the elasticity of capital supply (φ) and the supply elasticity of the other labor type ($\eta_{g'}$). Thus, when $\beta \leq 1$, the impact of a supply shock on native employment will be negative for skill group g if the weighted intensity of immigration in that skill group (first term in brackets) exceeds an appropriately weighted average of immigration intensity across all skill groups (second term in brackets).

In the standard case of a homogenous local labor supply elasticity (i.e., $\eta_U = \eta_S = \eta$), Equations (4) and (5) imply that both the wages *and* the employment of the skill group that experiences the larger migration-induced supply shock (i.e., the group for which $\pi_g^I/\pi_g^N > \Pi$) will decline relative to the wages and employment of the other group. These implications also hold for more general production functions than (1), such as functions that distinguish many skill groups (see, e.g., Dustmann, Frattini, and Preston 2013) or allow for a third nest within skill groups (see, e.g., Card and Lemieux 2001; Borjas 2003).

If, in contrast, the local labor supply elasticity varies across groups, then the wages of the skill group for which immigration is relatively intensive may *increase* relative to the other skill group, as can be shown by considering the relative wage effects:

$$dlog\mathrm{w}_S - dlog\mathrm{w}_U = \frac{(\beta-1)\left[\frac{\pi_S^I}{\pi_S^N}(1-\varphi\,\eta_U)-\frac{\pi_U^I}{\pi_U^N}(1-\varphi\,\eta_S)\right]}{1-(\beta-1)[\eta_S(1+s_S\phi)+\eta_U(1+s_U\phi)-\eta_U\eta_S\varphi]}dI. \tag{6}$$

Supposing that migration is predominantly unskilled (i.e., $\frac{\pi_S^I}{\pi_S^N} < \frac{\pi_U^I}{\pi_U^N}$) and that the local labor supply of the unskilled is elastic relative to that of the skilled (i.e., η_U is large relative to η_S), then the relative employment effect is amplified and the relative wage effect muted compared to the case of a homogenous local labor supply elasticity. Provided that capital is not fully elastic ($\varphi < 0$) and some skilled migrants enter the local labor market ($\pi_S^I > 0$), the wages of the unskilled may even increase relative to those of the skilled. At the same time, employment of the unskilled will strongly decline relative to that of skilled natives. Thus, in these "perverse" cases, relative wage and employment effects have the opposite signs. This observation emphasizes the need to investigate immigration-induced wage and employment responses *jointly* to avoid a misleading picture of immigration's overall labor market effects. It should further be noted that in the case of two skill groups, such an effect will only be observable when capital is not perfectly elastic; that is, $\varphi < 0$. If an additional skill group is added, perverse effects can occur even when the capital supply is fully elastic (see Online Appendix A.III).

II.D Wage Rigidities

Our analysis so far assumes that wages are fully flexible. However, in reality wages may, at least in the short run, be partially downward rigid, and the degree of wage rigidity may vary across skill groups (see, e.g., Card, Kramarz and Lemieux 1999). For instance, skilled workers may be more likely to be covered by long-term contracts than unskilled workers, preventing firms from immediately cutting skilled wages.[15] Next, we allow for partially rigid wages, and further allow the degree of wage rigidity to be different for skilled and unskilled workers.

Let $\overline{dlogw_g}$ denote the wage change, constrained by labor market institutions or private contractual arrangements, by which wages for skill group g may decline at most. The smaller (in absolute terms) $\overline{dlogw_g}$, the more rigid wages are. Provided that wages cannot fall by as much as the equilibrium wage response given by Equations (4) and (5) for both skill groups, the economy is demand-side constrained and there will be an abundance of native workers who would like to work for the current wage rate, but cannot find a job, and the employment response of natives is given by Equation (2) where wage responses dlogw_g are determined exogenously by the degree of wage rigidity $\overline{dlogw_g}$.

Heterogeneity in the degree of wage rigidity provides, in addition to heterogeneity in labor supply responses, an explanation for "perverse" effects in which the group that experiences the greatest shock needs not be the group that suffers the largest wage or employment decline.[16] For example, if wages of skilled workers are fully downward rigid whereas wages of unskilled workers are not, then employment differences between unskilled and skilled workers may be muted in response to a labor supply shock, compared to the case of a homogenous degree of wage rigidity (or the case of fully flexible wages and a homogenous labor supply elasticity). Furthermore, the

[15] Angrist and Kugler (2003) point out that labor market rigidities, while protecting some native workers from immigrant competition, can increase negative employment effects. They provide evidence that migration creates higher employment responses in countries with more rigid institutions.

[16] Wages of skilled workers are more downward rigid than those of unskilled workers if $\frac{\overline{dlogw_s}}{dlogw_s} < \frac{\overline{dlogw_u}}{dlogw_u}$, where $dlogw_g$ is the equilibrium wage response in the case of fully flexible wages given by Equations (4) and (5).

employment of skilled workers may decrease relative to unskilled workers even when more unskilled than skilled immigrants are entering the labor market.

III. Background and Data

III.A. Commuter Policy

Our analysis takes advantage of a commuting policy (*Grenzgängerregelung*), triggered by the fall of the Iron Curtain and implemented by the German government in 1991, that allowed workers from the neighboring Czech Republic to seek employment in German districts along the German-Czech border (see also Moritz, 2011 who was the first to investigate the labor market effects of that policy). Although allowed to work in Germany, these workers were not granted residence, forcing them to commute on a daily basis between their home country and their workplace in Germany, an aspect that our empirical analysis exploits (see Section IV.B.i).[17] The policy was otherwise nonrestrictive and unconstrained to specific industries or applicants with specific qualifications. Work permits were formally granted for up to two years and could be renewed after that.[18]

This particular commuting scheme was part of a larger scheme for the legal employment of foreign nationals in Germany announced in September 1990 and implemented on January 1, 1991, one year after the fall of the Berlin wall. The intention of the scheme's various provisions was to facilitate the recruitment of foreign workers in a time of increased labor demand following

[17] The requirement to commute was enforced via various channels. First, workers that entered employment under the commuting scheme had to apply for a special type of permit, the *Grenzgängerkarte,* which reflected the worker's conditional residence status. Second, in line with the German requirement that all residents register with the local registry office, a double registration was required by which both tenants and landlords had to submit information, making it impossible for Czech commuters to legally rent a home in Germany.

[18] Commuting requirements play also a central role in Angrist (1996) and Mansour (2010) who study the labor market response to exogenous changes in the commuting pattern of Palestinian day workers during the First and Second Intifada.

German reunification.[19] For example, a similar commuting scheme applied to Germany's second Eastern neighbor, Poland, and non-discriminately covered all German districts sharing a border with either Poland or the Czech Republic. The overall policy set up ensures that the commuting scheme examined here was exogenous to the economic conditions in the areas covered. We provide more details on the policy in Online Appendix B.

Figure 1 maps the region affected by the scheme, which comprises 21 districts within an approximate 80 kilometer band from the Czech-German border. Some of these districts, however, are close to the former East- and West German border and may thus have been affected after the 1990 reunification by commuters from East Germany, where wages were lower. Hence, to avoid any contamination of our experiment, we exclude districts located within approximately 80 kilometers of the former East- and West German border (although our results remain robust to less conservative choices). As Figure 1 shows, this exclusion leaves a rural region of 13 districts, or 291 municipalities, referred to hereafter as the "border region", which contains various small but no large cities. As column (1) of Table 1 illustrates, its local labor market at that time was characterized by a comparatively small share of highly skilled workers with university degrees, a young workforce, low wages, and a low share of preexisting immigrants.

The introduction of the commuting scheme in January 1991 led to a substantial and rapid inflow of Czech workers into the border region, whose employment shares in border and selected control districts (defined in Section IV.B) are plotted in Figure 2. By June 1992, the share of Czech nationals in the border region had increased from close to zero to about 3% and on average to about 10% in municipalities closest to the border. The employment share of Czech nationals in control districts, in contrast, being unaffected by the commuting scheme, remained negligible. As

[19] See "Anwerbestoppausnahme-Verordnung" (1990), Bundesgesetzblatt, Jahrgang 1990, Teil I.

Figure 2 also shows, the share of Czech workers remained stable from 1992 to 1993 and decreased thereafter, partly because of a stricter interpretation of the commuting scheme in later years, which was caused by allegations that the large Czech inflows into the border region had led to a worsening of conditions for native workers. Hence, in the empirical analysis, we focus on the immediate wage and employment effects of the labor supply shock up until 1993 because the "reverse experiment" of subsequent decline in the share of Czech nationals from 1994, albeit interesting, is potentially endogenous to local labor market conditions. We abstain in this paper from a detailed analysis of also long run outcomes.

Table 2 provides descriptive statistics for both the existing stock of workers in the border region in 1989 (i.e., before the entry of workers from the Czech Republic) and for Czech nationals, with their characteristics as of 1992. According to the table, Czech workers were far more likely to be unskilled (i.e., had no post-secondary degree) than the existing workforce (50.5% vs 27.6%) and more likely to fall into the 30 to 49 age group (61.9% vs 40.8%), with a much lower share of workers over 50 (3.7% vs 15.7%). The Czech nationals were also predominantly male, and in terms of concentration, overrepresented in construction, the hotel and restaurant industry, and wood processing and manufacturing, and underrepresented in the public sector. On average, Czech nationals earned 0.35 log points lower wages than natives. Interestingly, a significant wage gap between Czech and German nationals persists within detailed occupation groups (0.30) as well as within firms and occupation groups (0.22), pointing to the possibility that Czechs were willing to work for lower wages than Germans.

III.B. Data

Our data come from over two decades of German Social Security Records (from 1980 to 2001), which include all men and women covered by the social security system, excluding civil servants, the self-employed, and military personnel.[20] Three characteristics make this data set well suited for our analysis. First, the large sample size allows us to obtain fairly precise estimates of immigration on wages and employment even for detailed subgroups, although only a relatively small local area is affected by immigrant inflows. Second, the longitudinality of the data allows us to investigate whether the employment effects are driven by an increased outflow of workers into other areas or non- or unemployment, or by a decreased inflow of workers into the local labor market, a dynamic so far underexplored in the literature. Third, in addition to information on education, age, and other individual characteristics, the data include the citizenship of every employed individual, which allows identification of all Czech workers working in Germany but living in the Czech Republic. As a result, in our analysis, sampling error in the migration-induced supply shock, which attenuates the impact of the shock on native labor market outcomes (Aydemir and Borjas 2011), is close to zero.

Because our data set is constructed to observe each individual as of June 30 each year, each individual's employment status also refers to this date. The wage variable, in contrast, records the average daily wage in the employment spell that contains the reference date.[21] As is typically the case with social security data, our wage variable is right-censored at the social security limit, which in our sample affects only about 3% of all observations. Following Dustmann et al. (2009), we impute censored wages under the assumption that the error term is normally distributed while

[20] In 2001, 77.2% of all workers in the German economy were covered by social security and are hence recorded in the data (Federal Employment Agency, 2004).
[21] Because employers are required to update records only at the end of each year, this variable may also capture wage changes that occurred from June 30 to December of the same year.

allowing for a different residual variance by gender as well as by district. Information on districts or municipalities in our data refer to the individual's place of work and not her place of residence.

We distinguish two skill groups: *unskilled* workers who enter the labor market without postsecondary education and *skilled* workers who have completed an apprenticeship scheme or equivalent or graduated from a university or college.[22] This classification is particularly meaningful in the German context in which many apprenticeship jobs educate for professions that require college degrees in Anglo-Saxon countries (e.g., medical assistant or bank clerk).[23] We do not report separate results for university or college graduates because their share in the border region in 1990 was less than 5%. Since the Czech Republic runs, like Germany, a large scale apprenticeship system,[24] incoming Czech nationals fit easily into our skill classification scheme. Within each of these skill groups, we also distinguish three age groups: younger than 30, 30 to 49, and 50 and older. We further restrict the analysis to individuals aged between 18 and 65 and exclude irregular, marginal, and seasonal employment, as well as individuals undergoing apprenticeship training whose wages may not reflect their productivity. Our analysis of employment effects is thus based on regular full- and part-time workers, with part-time work (>30 hours per week) down-weighted into full-time equivalent units by 0.67 (18–30 hours) or 0.5 (<18 hours). Our wage analysis is based on full-time employees only.

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[22] To improve the consistency of the education variable in our data set, we impute missing values using past and future values of the education variable (see Fitzenberger, Osikominu, and Völter 2006). The imputed education variable is missing for 3.9% of observations in the overall data, and 2% of sampled observations in the border region. We classify these individuals as unskilled, although doing so has little impact on our findings.

[23] Such apprenticeship, which in Germany is highly structured, contains a school-based general education component, and takes between two and three years, explains why college or university enrolment in Germany is low in comparison to, for example, the U.S., being about 25% in recent cohorts and even lower in earlier cohorts.

[24] In Czechoslovakia, about 15% of those who finished compulsory education in 1989 entered the university track, compared to about 18% in Germany, while about 60% entered apprenticeships, compared to about 65% in Germany (OECD, Country Note on the Czech Republic, 1997).

IV. Empirical Strategy

In this section, we first explain how our main regression equations relate to the theoretical model presented in Section II and then describe our procedures for estimation and identification.

IV.A. Effect of Immigration on Wages and Employment

Our basic estimation equation regresses the change in log wages of natives (*N*) in skill group g, age group *s,* and area j between two periods, *t* and *k,* $\Delta lnw_{gs,j}$, or the percentage change in native local employment, $\Delta L_{gs,j}^{N}$, on the *total* inflow of Czech workers between 1990 and 1992 as a share of total employment in that area in 1990, ΔC_j^{92-90}:

$$\Delta lnw_{gs,j} = \alpha_{gs} + \beta_{gs}\Delta C_j^{92-90} + u_{gs,j} \quad (7)$$

and

$$\Delta L_{gs,j}^{N} = \gamma_{gs} + \delta_{gs}\Delta C_j^{92-90} + v_{gs,j}, \quad (8)$$

where

$$\Delta C_j^{92-90} = \frac{L_{j92}^{Czech} - L_{j90}^{Czech}}{L_{j90}^{N} + L_{j90}^{foreign}} \text{ and } \Delta L_{gs,j}^{N} = \frac{L_{gs,jt}^{N} - L_{gs,jk}^{N}}{L_{gs,jk}^{N}}.$$ [25]

Equations (7) and (8) are written in first differences to eliminate time-constant area and group fixed effects while allowing for time and skill group-specific growth rates in wages and employment, α_{gs} and γ_{gs}. The parameters of interest are β_{gs} and δ_{gs} which measure the impact of an inflow of Czech workers between 1990 and 1992 on the percentage change in wages and

[25] We scale the inflow of Czechs between 1990 and 1992 by *total* (including foreign) employment in 1990, as the supply shock may displace not only native but also pre-shock foreign workers. This choice has little consequences for our estimates as the share of foreign workers was small in 1990. It ensures that the coefficient δ_{gs} will be equal to -1 under full displacement, where every Czech worker displaces either a native or foreign resident worker in proportion to the employment share of each group. Similarly, because nearly all of the Czech inflow occurs in 1991 and 1992, the share of Czech workers in 1993 is almost the same as in 1992, such that adding another year into our measure of the supply shock barely changes our results.

employment of native workers in skill group gs in area j between the two time periods specified in Section IV.B.iv below. If wages are fully flexible, these parameters correspond to the expressions derived in Equations (4) and (5), Section II.B. Their sign and magnitude depend on the weighted difference between the density of immigrants in group gs ($\frac{\pi_g^I}{\pi_g^N}$) and the average density of immigrants across all other groups (Π), with the weights being a combination of the elasticity of local capital supply, and the gs group and other group's elasticity of local labor supply. If wages are partially rigid, the wage response β_{gs} is determined exogenously by the degree of rigidity (see Section II.D) and the employment response δ_{gs} is given by Equation (2), Section II.A.

This empirical specification not only corresponds to our theoretical setup but has the added advantage that identification of β_{gs} and δ_{gs} does not require the pre-allocation of immigrants to skill groups based on their observable characteristics. It thus avoids the problem of misclassification that arises when such observable characteristics are used to assign immigrants into skill groups in which they do not compete with natives.[26] Hence, not only is this specification consistent with our experiment—because only the total inflow of Czechs into the border region can be considered quasi-random—but the estimated parameter is clearly policy relevant in that it captures the total effect of the aggregate supply shock for specific groups of natives.

It should also be noted that the employment regression in Equation (8) differs from regressions typically estimated in the literature in two ways, which may lead us to detect a larger employment response. First, whereas we use the percentage change in local native employment ($\Delta L_{gs,j}^N$) as the dependent variable, many extant studies use the change in the (local) employment-

[26] Dustmann and Preston (2012) illustrate that assigning immigrants to skill groups based on observed characteristics may lead to serious misclassification because immigrants often *downgrade* upon arrival.

to-population ratio instead (see, e.g., Altonji and Card 1991; Dustmann, Fabbri, and Preston 2005; Boustan, Platt, Fishback and Kantor 2010, Smith 2012).[27] This alternative measure may be problematic: if the local native population adjusts in response to an inflow of immigrants, it would lead to smaller effect in the employment ratio, something we demonstrate in Section V.D.ii. Second, whereas we measure the labor supply shock as the ratio of employed Czechs and employment in the base period, the two most commonly used measures of a labor supply shock in the extant literature are the change in the total immigrant share (employed or not) in the population (e.g., Altonji and Card 1991; Dustmann, Fabri and Preston 2005; Dustmann, Frattini and Preston 2013) and the ratio of all incoming immigrants and the native (or resident) population in the base period (e.g., Card 2001, Card 2009, Boustan, Platt, Fishback and Kantor 2010). Like our measure, the latter (but not the former) avoids confounding the (exogenous) inflow of immigrants with a possibly endogenous adjustment in native population. However, unlike in our specification, the slope coefficient in this latter specification will—if the employment rates of recent immigrants and natives differ—be different from -1 even if every immigrant *who finds a job* displaces a native worker. Our specification, in contrast, measures the extent to which immigrants crowd out native employment irrespective of immigrants' willingness or ability to find a job.[28]

[27] An exception is Wagner (2010) who considers percentage changes in employment at the local labor market and industry level.

[28] This may be advantageous because the employment-to-population and participation rates of immigrants vary across countries, immigrant groups within countries, and time. For example, Card (2001) reports that in 1990 the employment rate was significantly *lower* for recent immigrants than for U.S. natives, while the OECD reports that in 2007 the rate has been *higher* for recent immigrants in both the U.S. and other countries (see Table II.3 in the International Migration Outlook 2009, OECD).

IV.B. Estimation and Identification

IV.B.i. Exploiting Distance to Border

The quasi-natural experiment of suddenly and unexpectedly opening border regions to commuters from the Czech Republic allows us to identify the effect of the induced labor supply shock on native workers' wages (Equation (7)) and employment (Equation (8)) in various ways. One option would be to compare the entire border region eligible under the commuting policy with suitable control areas that were similar in observable characteristics but not eligible. However, the nature of the commuting experiment provides additional variation in the exposure of different areas to Czech inflows that can be usefully exploited: because Czech workers were forced to commute daily, increased traveling costs exposed municipalities close to the border more to the policy. In fact, as demonstrated in Section V.A, distance to the border was a key determinant of where Czech workers located within the border region, explaining 38.7% of the overall variation in Czech employment across municipalities (see also Figure 3 and Table 4). We could therefore also estimate Equations (7) and (8) only for municipalities within the affected border region, using distance from the border region as an instrument. In our baseline specification, we combine the two approaches by pooling municipalities in the border region with unexposed control districts, thus exploiting variation in the employment share of Czechs within the border region in addition to using areas further inland as control units. To test the robustness of our findings we also report separate estimates based on the other two approaches, showing that all three approaches produce similar results.

IV.B.ii. Assumptions

For distance to border to be a valid instrument, the following assumptions need to hold. First, and most important, in the absence of a Czech inflow, the evolution of subgroup-specific local wages and employment must be uncorrelated with distance from the border. We provide support for this assumption in Table 3, by analyzing whether prior to the introduction of the commuting policies, municipalities in the border region closer to the border experienced differential trends in subgroup-specific outcomes from municipalities further away from the border. Reassuringly, the table shows that distance to border is, with one exception, uncorrelated with pre-policy trends in outcomes. Nevertheless, to make sure that our results are not driven by differential pre-existing trends, we report results with and without controls for municipality-specific time trends. We further estimate placebo regressions in pre-policy periods and adopt an event study approach (for some outcomes) to illustrate graphically that distance to border affects native local wage and employment growth only after the inflow of Czech workers actually occurred (Section *IV.B.iv*).

In addition, for β_{gs} and δ_{gs} in Equations (7) and (8) to correspond to their theoretical counterparts in Equations (4) and (5), "control" areas—that is, municipalities in matched inland districts and municipalities at the edge of the border region that received barely any Czech commuters—must not be affected by the Czech inflow into "treated" areas near the border. This condition would be violated if natives from the treated municipalities moved to control areas in response to the Czech inflow, thereby increasing employment and lowering wages in these areas. Because the labor supply shock to the border region was negligible in national terms, matched control areas away from the border region are clearly unaffected by this shock. However, municipalities at the edge of the border region could in principle be affected by a much larger Czech inflow to municipalities right on the border if natives from these municipalities responded to the labor supply shock by moving to nearby municipalities within the border region that received fewer Czech commuters. In this case, the employment response from our baseline specification might simply reflect employment shifts across municipalities within the border region that left overall employment in the border region unchanged. We investigate this possibility by comparing the region very close to the Czech-German border with a set of control districts located sufficiently far from the German-Czech border, thereby discarding any variation in Czech inflow within the border region. We implement such a comparison using a synthetic control method (detailed in Online Appendix C) that compares a single treatment to a weighted average of available control units (see Abadie, Diamond, and Hainmueller 2010). This approach yields wage and employment effects that are broadly similar to our baseline specification.

Finally, we need to rule out that the opening of the Czech-West German border directly affected areas close to the border, other than through the increased inflow of Czech workers into those areas. One channel through which the opening of the border could affect areas close to the border is increased trade between the border region and the Czech Republic or increased foreign direct investment (FDI) by firms in the border region in the Czech Republic. An alternative channel could be increased market access: areas close to the border may benefit from the opening of the border by occupying a more central position within Germany and Europe. We believe that both channels are unlikely, for two main reasons. First, in 1993 (the last year in our main empirical analysis), some trade restrictions between Germany and the Czech Republic were still in place and the trade volume between the two countries was still very low. Similarly, throughout the mid-1990s, German FDI in the Czech Republic was relatively small in magnitude and concentrated in the capital Prague, rather than in areas close to the German border. In addition, as shown by Redding and Sturm (2008), gains from trade take a long time to materialize, whereas we focus on short-term effects in the immediate aftermath of the border opening. Second, such shocks, if present, would be likely to affect the border region as a whole, but when we drop control districts further inland from our sample and exploit variation in Czech inflows within the border region only, our estimates are very similar.[29]

IV.B.iii. Selecting Control Areas

For the matching of control areas, we consider only West German districts of similar urban density (rural areas or areas with intermediate agglomerations) and exclude districts within 80 kilometers of the former border between East- and West Germany to avoid contamination from German reunification. Matching is based on variance-weighted differences in the employment share of the education groups, the employment share of foreign nationals, mean log wages, the share of right-censored wage observations, local employment levels, and the employment shares of four age groups in 1989 (the year before reunification and the fall of the Iron Curtain). What is most important is that in these calculations, we do *not* match on preexisting time trends. The 24 matched control districts (corresponding to 1259 control municipalities) depicted in Figure 1 are much more similar to border districts than other West-German districts (see column (3), Table 1).[30] Our baseline specification thus refers to 1550 municipalities (291 in the border region and 1259 control municipalities). However, the exact number varies slightly across subgroups and years, as there are some small municipalities which do not employ workers of a specific type or in a specific year.

[29] It is in principle possible that even within the border region, municipalities close to the border suffer a negative demand shock relative to municipalities further away from the border when the border opens, because Germans start buying goods in the Czech Republic. This is unlikely during the period we consider, as the Czech Republic was separated by the iron curtain for more than four decades, and products from its formerly state-directed economy were of little appeal to Western consumers. To test for this nevertheless, we re-estimated our model excluding the retail sector. We find very similar employment- and wage effects than those reported.

[30] Three out of the 24 matched control districts are located next to the border region and thus could potentially be affected by spillover effects from the border region, caused by workers moving from the border region to these districts. Our findings are very similar if we exclude neighboring control districts from the sample.

When using the synthetic control approach, in contrast, which discards all variation across municipalities within the border region, we match similarly on the education, foreign and age shares, but also on outcome variables from 1986 to 1989 (cf. Abadie, Diamond, and Hainmueller 2010). In these estimates, therefore, we explicitly match on preexisting time trends (see Online Appendix C for details).

IV.B.iv. Timing and Placebo Tests

When estimating Equations (7) and (8), we adopt a flexible specification that allows us to assess how quickly local wages and employment adjust in response to the labor supply shock. Although the regressor ΔC_j^{92-90} is always defined as the inflow of Czech workers into area *j* between 1990 and 1992 as a share of local employment, we specify the dependent variables over different time windows. In particular, we obtain coefficients from annual regressions of wage or employment changes between the years *t* and *t-1* on ΔC_j^{92-90}, instrumented with distance to border. To obtain the overall impact of the labor supply shock over longer periods, we then sum the respective coefficient estimates for t = 1991 to t = 1993.[31] Running yearly rather than long difference regressions is not only informative about the timing of adjustment but allows us to address potential selectivity bias in wage estimates, as the employment response to a labor supply shock may differ across the wage distribution (see Bratsberg and Raaum, 2012, and Llull, 2013, for a discussion). To deal with selection, we restrict in the wage analysis the sample to individuals who are employed in the municipality in both *t* and *t-1*, thus keeping the composition of workers

[31] We implement the IV estimator in two steps: a first stage estimation at the municipality level, regressing ΔC_j^{92-90} on distance to border and its square and weighting each observation by total employment in the municipality in 1990; and a second stage regression of subgroup-specific native employment and wages in the municipality on the predicted inflow of Czechs, $\Delta \hat{C}_j^{92-90}$, with each observation weighted by subgroup-specific employment in t-1.

constant over the two time periods. As illustrated below, we find that if instead longer differences are estimated on data that discard longitudinal worker information, selective employment response does indeed lead to underestimation of the wage effects.

We also estimate Equations (7) and (8) for the years prior to 1990, when the later inflow of Czech nationals should have no impact on native employment changes. Formulating the hypotheses H_0: $\beta_{gs} = 0$ and H_0: $\delta_{gs} = 0$ for $t \leq 1990$ provides a placebo setup against which to probe the identifying assumption that areas located close to the border experienced the same time trends prior to 1990 as areas located further away. Since we did not match on preexisting trends when selecting control districts, these tests provide a valid falsification test.[32]

Since our estimation strategy proceeds in multiple stages, which makes the computation of analytical standard errors complicated, we bootstrap standard errors using the wild-bootstrap procedure and 500 repetitions (see Cameron, Gelbach and Miller 2008). While our analysis is performed at the municipality level, we cluster standard errors at the district level. For our main outcomes of aggregate and skill-specific local wage and employment effects of natives, we additionally report standard errors based on the Spatial Heteroscedasticity and Autocorrelation Consistent (SHAC) variance estimator proposed by Conley (1999) and adopted by for example Kline and Moretti (2014), which allows for correlation between areas that are geographically close but belong to different administrative units (see column (6) in Table 6).[33]

[32] Angrist and Krueger (1999) implement a similar test, illustrating that the estimated effect of the Mariel Boatlift on the Miami labor market is sensitive to differences in trends between treatment and control units.

[33] There are various difficulties in applying this procedure to our context since, in contrast to Kline and Moretti (2014), our estimation strategy proceeds in multiple stages. We have implemented the SHAC standard errors in our long difference specification, ignoring uncertainty from the first stage. Additional robustness checks show that ignoring the uncertainty in the first stage has little impact on the standard errors (see Table O.2 in the Online Appendix).

V. Results

V.A. First Stage: Distance to Border and Location of Czechs

In Figure 3, we plot Czech employment growth from 1990 to 1992, ΔC_j^{92-90}, in municipalities within the border region against the municipality's distance to the closest border crossing, $\mathrm{Z_j}$, weighting municipalities according to 1990 employment levels. As the figure illustrates, distance to border is indeed a key determinant of where Czech nationals located within the border region: municipalities next to the border received the largest inflow of Czech workers, corresponding on average to almost 10% of employment in 1990. Municipalities located more than 50 kilometers away from the border, in contrast, experienced hardly any inflow.

We report the corresponding regression results (the first stage), approximating the relationship between the inflow of Czech nationals and distance to border as a quadratic function, in Table 4, reporting results for the border region only in column (2) and for the estimation sample including matched control districts in column (3). The coefficients on distance and distance squared are jointly highly significant (F = 42.58) and together explain 38.7% of the variation in the Czech employment share across municipalities within the eligible border region or 54.4% of the variation across border and matched control municipalities.[34]

[34] Analysis by skill group further reveals a similar relation between distance to border and Czech inflow for skilled and unskilled workers, and thus no significant association exist between distance to border and the unskilled-skilled ratio of Czech commuters. We have also estimated a variety of alternative first stages based on different functional form assumptions (i.e., a third order polynomial and a spline function in distance to border) and different distance measures (driving distance and driving time). These alternative specifications yield very similar first stage and 2SLS estimates (see Table O.1 in the Online Appendix D).

V.B. The Impact of Czech Inflows on Native Wages and Employment

V.B.i. Aggregate Wage and Employment Effects

Figures 4a and 4b provide a first visual assessment of the Czech inflow's effect on the local wages and employment of all native workers in the municipality. These figures are based on our estimations of Equations (7) and (8), which regress municipality-level changes in native wages or employment between two consecutive years on the Czech inflow between 1990 and 1992 (except for 1991, which is based on the 1990–1991 inflow) instrumented by the municipality's distance to the border. We then plot the cumulative effects relative to 1990 by summing the estimated slope coefficients backward and forward. The outcomes thus represent the cumulative wage (employment) effects of the Czech inflow between 1990 and 1992 for each year between 1986 and 1995. We display the corresponding cumulative post-policy regression coefficient in 1993 in row (i) of Table 5 (Panel A).

As the figure shows, prior to 1990, the estimated coefficients for both employment and wages are small and statistically not significantly different from zero, meaning that distance to border does not help to predict local employment and wage trajectories in the pre-policy period (see also Table 3). We thus cannot reject the falsification test described in Section IV.B.iv. After the policy comes into effect in 1990, however, local wages—and in particular local employment of native workers—drop significantly. Whereas wages respond immediately, the employment effect builds up and employment continues to decline from 1992 to 1993, although the employment share of Czech workers reaches its peak in 1992. By 1993, a 1 percentage point increase in the inflow of Czech workers relative to native employment in the baseline has led to about a 0.13 percent decrease in native wages, a 0.93 percent decrease in native local employment, and a 0.07 (1-0.93) percent increase in total (including Czech) local employment. Putting the wage response into perspective, the real wage growth over the period considered of workers employed in the two consecutive periods was about 3 percent per year, meaning that the negative wage effects do not necessarily imply a decline in natives' real wages.

Interpreted within the simple model laid out in Section II, these negative overall wage and employment effects suggest that at least in the short run, the local supply of capital is not fully elastic. The large employment response, coupled with a smaller wage response, could either be driven by a high local labor supply elasticity or by wages being partially downward rigid in the short run, or both. Our estimates further imply a wage elasticity ($\varphi = -\frac{\alpha\lambda}{1-\alpha+\lambda}$ in our model) of 0.54 (0.07/0.13), which is well within the range of existing estimates ranging from 0.15 to 0.75 (e.g., Hamermesh, 1993; Lichter, Peichl and Siegloch, 2015).[35]

V.B.ii. Wage and Employment Effects by Skill Group

According to Table 2, a higher fraction of Czech commuters was unskilled relative to natives. We would therefore expect the overall inflow of Czechs to depress local wages and employment of native unskilled workers by more than those of native skilled workers, unless the two groups differ sharply in their local labor supply elasticity or the degree of wage rigidity; see Equations (2), (4) and (5). We indeed find that both the wages and employment of native unskilled workers do decline relative to native skilled workers (see Table 5, Panels B and C). Over the 1990–1993 period, a 1 percentage point increase in the employment share of Czech workers decreases the local wages and employment of unskilled natives by 0.2 and 1.37 percent, respectively, but of

[35] As argued by Borjas (1999, 2013), the simple model laid out in Section 2 with a Cobb-Douglas production function implies, in the case of fully inelastic capital, a wage elasticity equal to (minus) the labor's share of income, which is typically estimated to be around 0.3. Our estimate is reasonably close, especially when taking into account the noise in the wage and employment estimates.

skilled natives by only 0.11 and 0.50 percent. Rows (ii) of Table 5 reports simple OLS estimates that do not instrument the share of Czech workers by distance to the border. Here, the estimated wage and employment effects are smaller than the IV estimates, particularly for skilled workers. This outcome is to be expected if Czech workers predominantly entered municipalities experiencing higher employment and wage growth.

As an alternative measure for skill, we use individuals' occupation (see Figure 5). Specifically, we estimate our baseline specification separately for nine 1-digit occupations and plot the resulting 2SLS cumulative (1990 to 1993) wage (Panel A) and employment (Panel B) coefficients in an occupation against the occupation's exposure to the labor supply shock, measured as the employment share of Czech commuters in the occupation in 1992 divided by the average share. The figure clearly demonstrates that, in line with standard models of immigration, local wages and employment of natives declined more in occupations with a larger exposure to Czech workers.

V.C. Robustness Checks

V.C.i. Common Time Trend

The findings in Table 5 are robust to a number of specification checks, reported in Table 6 and using the 2SLS estimates from Table 5 (row (i)), as a reference point. In column (2), we account for possible municipality-specific time trends, identified based on 1987–1989 data, and report trend-adjusted estimates. The employment estimates are similar to those reported in column (1), while the wage estimates are larger in magnitude. In column (3), we drop all control districts and compare only differentially exposed areas within the border region, whereas in column (4) we compare the region very close to the Czech German border—which we refer to from now on as the "inner" border region" and which, because of shorter distance to the border, received the vast majority of Czech inflows—with unaffected control areas.[36] In both cases, the results are very similar to those for our baseline estimates, indicating not only that our findings are not dependent on the particular matching of control districts, but providing also indirect support for our identifying assumption that in the absence of a Czech inflow, the evolution of subgroup-specific local wages and employment is likely to be uncorrelated with distance from the border.

V.C.ii. Worker Selection

Column (5) of Table 6 reports the results of estimating Equations (7) and (8) in long differences; that is, regressing local wage and employment growth *between 1990 and 1993* on Czech inflows between 1990 and 1992 rather than estimating *annual* regressions and summing the coefficients as in column (1). In these calculations, as is common in studies using repeated cross-sectional data, (log) wages are averaged over *all* workers who are employed in any of the two years, 1990 and 1993, rather than over workers who remain employed in the district in two consecutive years as in our baseline specification. As expected, the employment effect estimates are barely affected and remain very similar to those in the first column.[37]

These calculations do, however, highlight the importance of how wages are measured. The results of the long difference estimations point to no significant wage effects of the Czech inflow for either skill group. This finding suggests that the workforce composition changes as a result of

[36] We split municipalities within the border region according to their fitted values from the first stage regression. The inner border region is comprised of 145 municipalities in which the predicted inflow of Czech was above the median, averaging to about 5.8% of total employment.
[37] The small difference arises for two reasons: First, the baseline specification weights the annual regressions by group-specific employment in *t*-1, which changes slightly from year to year, while the long difference regression references only 1990. Second, the sum of annual employment growth rates ($\frac{L_{jt}^N - L_{jt-1}^N}{L_{jt-1}^N}$) does not correspond exactly to the employment growth rate over three time periods ($\frac{L_{jt}^N - L_{jt-3}^N}{L_{jt-3}^N}$).

the labor supply shock, with low-wage workers more likely to leave or not enter the workforce in response to migration. Hence, a simple comparison of average wages before and after the migration-induced supply shock underestimates the wage effect on the remaining workers, meaning that if immigration leads to selective employment effects, estimations based on repeated cross sections some years apart may underestimate, or even fail to detect, adverse wage effects.

For this long-difference specification, we additionally report standard errors based on Spatial Heteroscedasticity and Autocorrelation Consistent (SHAC) variance estimator proposed by Conley (1999) (column (6)), which allows for correlation between areas that are geographically close but belong to different districts. Standard errors are very similar to our baseline bootstrapped standard errors which allow for clustering at the district level.

V.C.iii. Synthetic Control Approach

An alternative estimation strategy, the synthetic control approach, discards the variation in municipalities' exposure to Czech commuters induced by distance to the Czech-German border, and instead compares wages and employment of the entire (inner) border region with those in the matched control districts. It thus internalizes all employment movements across municipalities within the border region. To obtain sharper results, we drop the outer border region from the sample and instead compare the evolution of aggregate native employment and wages in the highly exposed inner border region (treatment unit) with that in unexposed control districts (synthetic control units). Figure 6 display the evolution of the native wage (Panel A) and employment (Panel B) gaps between the inner border region and its synthetic control (bold line). As the figure shows, whereas both native wages and employment unfold in almost the same way in the treatment and control units prior to the policy (recalling that in contrast to Figures 4a and 4b, we are now explicitly matching on *trends*), in 1991 a gap begins to emerge in the treatment area relative to the control areas. To assess the statistical significance of this divergence, the figure also displays permutation tests in which we apply the synthetic control method to every potential control in our sample (as in Abadie, Diamond and Hainmueller, 2010). The results show that the employment but not the wage gap is exceptionally large in the treated inner border region compared to placebo districts, indicating statistical significance of the employment but not the wage gap. It is not surprising that outcomes from the synthetic control approach are more noisily estimated than our baseline estimates, as this approach discards any variation in the inflow of Czech workers within the inner border region.

To compare these outcomes with our estimates for the impact of the inflow of Czech workers on native local wage and employment growth, we must scale the wage and employment gaps (-0.007 and -0.079 by 1993) by the share of Czech workers that entered the treatment region (5.8%). The results, -0.12 for wages, and -1.36 for employment, are roughly in the same ballpark as our baseline coefficients of -0.134 and -0.926 in row (i) of Table 5 (Panel A).[38] Hence, the inflow of Czech workers into the inner border area led to an overall decline in native employment in that region and not merely to employment shifts across municipalities within the region.

V.D. Age Group-Specific Responses

In Table 7, we provide a more detailed analysis by investigating whether the Czech inflow affects labor market outcomes differently for younger (<30) and older (≥50) natives. The estimates refer to our baseline specification, which links the *overall* inflow of Czech workers to skill- and age-

[38] Since the synthetic control approach explicitly matches on pre-trends, it may be more appropriate to compare results from this approach with our estimates that adjust for pre-trends presented in column (2) of Table 6. Trend-adjusted estimates are with -0.20 and -0.927 similar to the unadjusted estimates.

specific wage and employment growth (Equations (7) and (8)) and captures the cumulative effects up until 1993. We report two types of estimates: those that are not trend adjusted (columns (1), (2)) and those that are (columns (3), (4)).

The findings point toward perverse effects across age groups, in particular for skilled workers. As Table 2 shows, in comparison to natives, Czech workers were more concentrated in the medium age range (30–49) than among young and older workers. Thus, according to standard immigration models, which restrict labor supply elasticities (or the degree of wage rigidity) to be the same across age groups (and may allow for imperfect substitutions), both employment and wages should decline the most for the middle-aged within each skill group. The estimates in Table 7, however, suggest that among skilled workers, young workers below 30 suffer the largest wage loss of all three age groups, whereas older workers aged 50 and above suffer the largest employment losses. The pattern is similar among unskilled workers.

Our model provides two complementary explanations for these patterns. First, older workers may be more elastic in their labor supply than younger workers. This is plausible, as older workers may have easier access to social security and unemployment benefits than young workers—who in turn may be willing to accept wage cuts at the beginning of their career to avoid scaring.[39] According to this explanation, we would expect non-employment movements to be relatively more important for older than young workers, a conjecture we confirm below. Second, wages may be more downward rigid for older than for younger workers as older workers may have more permanent and rigid contracts than younger workers. Also, younger workers are typically on a steep wage growth path whereas wages of older workers no longer grow (recall that our wage analysis is restricted to workers who are employed in two consecutive time periods). Thus, real

[39] See Ruhm (1991) for a discussion. Scarring effects can be due to a loss of human capital, which is particularly costly for young workers, signaling effects (Lockwood 1991), or discouragement and habituation (Clark et al. 2001).

wage cuts are particularly visible and difficult to implement for older workers, while for younger workers they imply lower growth than they otherwise would have.

V.E. Margins of Adjustment

V.E.i. Inflows versus Outflows

The overall employment effect reported above can be decomposed into workers who leave employment in a particular area (outflows) and workers who do not enter employment but would have done so in the absence of the labor supply shock (inflows), i.e., $\frac{L_{gt}^N - L_{gt-1}^N}{L_{gt-1}^N} = \frac{\text{Inflow}_g^N}{L_{gt-1}^N} - \frac{\text{Outflow}_g^N}{L_{gt-1}^N}$. In Figure 7 we report estimates of the labor supply shock-induced impact on the inflow (Panel A) and outflow (Panel B) rates using the same regressions as in Figure 4, with the overall inflow and outflow rates as the dependent variable. Unlike Figure 4, however, Figure 7 represents yearly rather than cumulative responses. The results indicate that overall yearly employment effects are driven primarily by a reduction in inflows and to a far lesser extent by an increase in outflows. Moreover, whereas the inflow response is immediate, the outflows response is delayed and begins increasing only in 1991, one year after the policy came into effect. This immediate response to inflows helps explain why native employment levels seem to react so rapidly to local shocks.

Table 8 provides detailed estimates using column (1) as the reference for employment effects in the aggregate (Panel A) and for different skill groups (Panel B) and age groups (Panel C).[40] Columns (2) and (3) show the reduction in inflows and increase in outflows, respectively,

[40] Because inflow and outflow rates tend to be smaller in municipalities close to the border, the overall employment estimates reported in Table 8, column (1), are trend adjusted and differ slightly from our baseline estimates in Table 5. Since a large NATO cold war military exercise (REFORGER 88*)* in the border region in 1988 coincided with an unusually large outflow of workers from the 1987–1988 social security records who returned in 1989, we use only

each of which makes up roughly 17% of average employment over the 1985–1989 pre-policy period. All table entries refer to the overall effect by 1993, and, as before, are obtained by summing the coefficients from the annual regressions. As Figure 7 and the table entries indicate, inflows are far more important than outflows for explaining the total employment response in all skill and age groups, accounting for at least 87% of the overall reduction in employment, with the exception of older workers, for whom outflows make up 28% of the overall employment effect.

This finding puts a new spin on the usual interpretation of employment responses to labor supply shocks. In particular, rather than implying that native workers lose their jobs as Czechs enter the local labor market, the large employment response is induced by workers not employed in the affected area (but possibly in other areas) at the time of the policy no longer being hired. One explanation for why the local employment decline is almost entirely absorbed by reduced hiring as opposed to increased separations is that this group of "outsiders" is particularly elastic in their labor supply because they have outside alternatives (e.g., the ability to move to an unaffected area). An alternative explanation (and one in line with our arguments in Section II.D) is that, because of private contractual arrangements or labor market regulations, it is costly for firms to lay off their existing workforce. Firms can, however, immediately adjust their hiring behavior. Whichever the explanation, most of the burden of the employment effect is borne by outsiders, whose strong labor supply response shields employed or incumbent workers from the labor supply shock.[41]

1989 and 1990 to account for municipality-specific preexisting time trends rather than the 1987–1989 period used in Table 6, column (2), and Table 7, columns (3) to (4).

[41] A related point is made by Cadena and Kovak (2013).

V.E.ii. Non-employment versus geographic movement

Another way to decompose the overall employment effect is to distinguish between movements from and to non-employment (including movements from and to unemployment) and movements from and to employment in other areas.[42] For simplicity, we refer to the latter as "geographical movements", but we would like to point out that in our data we observe whether between two periods an individual is employed in a different area, and not whether she actually moved to the area.[43] We report the decomposition in columns (4) and (5) of Table 8, defining a transition as a geographic movement if the worker is employed in one municipality in the base period and in another municipality one year later. In terms of magnitude, movement from and to non-employment is far more relevant than movement across areas: only about 17% (-0.168/-0.989) of the overall employment effect results from direct employment movement to and from other areas, with the remaining 83% stemming from movement into and out of employment (see Panel A). These entries, however, consider only direct movement in which the worker was employed in two consecutive years in different municipalities. But workers might decide to move away only after a spell of non-employment, meaning that the numbers in column (5) could underestimate the extent of geographic mobility. To investigate this possibility, we categorize inflows from and outflows to non-employment as geographic movement if *within three years* (as opposed to the previous or next year, as in column (5)) the worker is observed working in a different municipality.[44] This redefinition increases the estimated importance of geographic movement, which now rises from

[42] That is, $\frac{L_{gt}^N - L_{gt-1}^N}{L_{gt-1}^N} = \underbrace{\frac{\mathrm{In}_g^E - \mathrm{Out}_g^E}{L_{gt-1}^N}}_{\text{geographical}} + \underbrace{\frac{\mathrm{In}_g^N - \mathrm{Out}_g^N}{L_{gt-1}^N}}_{\text{non-employment}}$, where In_g^E are inflows from employment in other areas, In_g^N are inflows from non-employment, Out_g^E are outflows into employment in other areas and Out_g^N are outflows into non-employment.

[43] That is, not all of our geographical movements, and some of our non-employment movements, may in fact entail a change in residency.

[44] Even if we increase the time window to five years, the numbers are similar.

17% (column (5)) to close to 29% (column (6)) of the overall employment effect. Interestingly, a further decomposition of geographic movement into inflows and outflows shows that most geographic movement is driven by a reduction in *inflows*, with outflows to other areas being negligible (see Table 9). Thus, while we do find geographical responses to the labor supply shock, our results also indicate that these are not induced by individuals in affected areas seeking employment in unaffected areas (as usually suggested in the literature, see Peri and Sparber 2011 for a review), but by individuals not seeking employment in affected areas. Spatial arbitrage in response to local shocks may thus be achieved through a reduction rather than an increase in geographical mobility (if individuals who would have found employment in the affected area if the shock had not happened stay in their area instead). This pattern is consistent with recent evidence on internal migration rates in the U.S. during the Great Recession (see Monras, 2015).

Finally, in column (7) of Table 8, we directly investigate how the municipality's *population* responds to the Czech inflow using population counts from Germany's Federal Statistical Office, an alternative data set for assessing the magnitude of geographic movement, albeit one not broken down by age or skill. The results indicate that a 1 percentage point increase in the employment share of Czech commuters—who do not live in the affected German border region—decreases local population levels by 0.3%, a coefficient close to the geographic displacement of native workers estimated based on individual worker tracking in the social security data (column (6)).

Overall, these findings underscore that although non-employment movement clearly dominates, population does adjust in response to a labor supply shock even in the short run. This observation has important implications for the estimation of the employment effects, which would have been roughly 30% smaller if, as in much of the extant literature, we had measured them as changes in the employment-to-population ratio.[45]

V.E.iii. Geographic and Non-employment Movement by Skill and Age

The relative importance of geographic movements differs markedly between skill and age groups. Whereas for unskilled workers, the entire employment effect is due to movement into and out of employment, for skilled workers, between 25% and 37% of the overall employment response involves movement across areas, mirroring the larger geographical mobility rates for skilled than unskilled workers observed in our (see Panel B of Table 8) and other data (see for example Amior, 2015). Differencing between age groups in Panel C, direct and indirect (through non-employment) geographic movements are relatively more important for workers under 30 who experience the smallest absolute employment effect (27%, -0.147/-0.555, columns (4) and (5)). For workers over 50, in contrast, for whom the absolute employment effect is largest, nearly all the labor supply shock is absorbed through transitions into and out of non-employment—as we would expect if older workers are entitled to generous unemployment benefits and can take advantage of early retirement packages.

VI. Discussion and Conclusions

Exploiting a rare commuting policy that created a sharp, sudden, and unexpected inflow of Czech workers to areas along the German-Czech border, we assess the impact of an immigration-induced labor supply shock on native wages and employment and identify response dynamics in the

[45] When using the change in employment-to-population ratio in the municipality as our dependent variable, we obtain a coefficient of –0.611, with a standard error of (0.217).

policy's immediate aftermath. Mirroring the adjustment responses to economic downturns, our results show that the labor supply shock led to a moderate decline in local wages and a sharp decline in local employment—an effect that is nearly entirely accounted for by a reduction in hiring, and not by an increase in separations.

What, then, do our findings imply for the possible welfare consequences of the immigration-induced labor supply shock? Most important, the crowding out of natives need not signal that the immigration-induced labor supply shock caused affected workers to suffer large utility losses. In fact, the relatively modest wage and employment declines for natives who were employed when Czech workers entered the labor market ("insiders") suggest that their utility decreased little.[46] If interpreted through a fully competitive model with flexible wages, natives who were not employed (in the affected area) at the time of the shock ("outsiders") will not suffer large utility losses either. Such a model attributes the large employment response of outsiders to a large local labor supply elasticity—which must mean that these workers had been close to indifferent between employment in the affected area and other options; otherwise a small change in the (local) wage would not have made them change their employment status.

However, the large employment response of native outsiders may also be driven by labor market regulations or private contractual agreements that prevent firms from cutting wages and laying off its incumbent workforce, at least in the short run. If so, we would expect a decline in the utility of non-employed workers unable to find a job because of increased competition from the Czech workers. In this case, the welfare consequences of an immigration-induced labor supply shock would be unequally distributed even within detailed skill and age groups: its adverse effects

[46] Our baseline estimate suggests that even in the municipalities closest to the border (and thus most exposed to the labor supply shock), the wages of native incumbent workers employed in two consecutive years grew about 7.7% between 1990 and 1993 compared to 9% had the shock not happened.

would, at least in the short run, be borne predominantly by non-employed outsiders, whereas institutional constraints insulate insiders from the increased competition.

It is important to emphasize that we focus on the *short-term* effects of an immigration-induced labor supply shock, which may be more pronounced than the longer-term effects typically considered in the literature. For instance, wages may be partially downward rigid in the short but not in the longer term, while the supply of capital may be more responsive—especially if, as in our case, the inflow of immigrants was unexpected. In the longer term, firms and workers could also respond to an immigration-induced labor supply shock along other dimensions not considered here. For instance, firms might change their technology (see, e.g., Lewis 2011), labor market entrants might invest more in full-time education (see e.g., Smith 2012 and Hunt 2012), and even experienced workers might upgrade to more skilled occupations (Peri and Sparber 2009). However, the short-term responses that we investigate here shape agents' incentives to undertake such investments, and are therefore crucial to understand and assess the mechanisms behind any longer-term adjustment.

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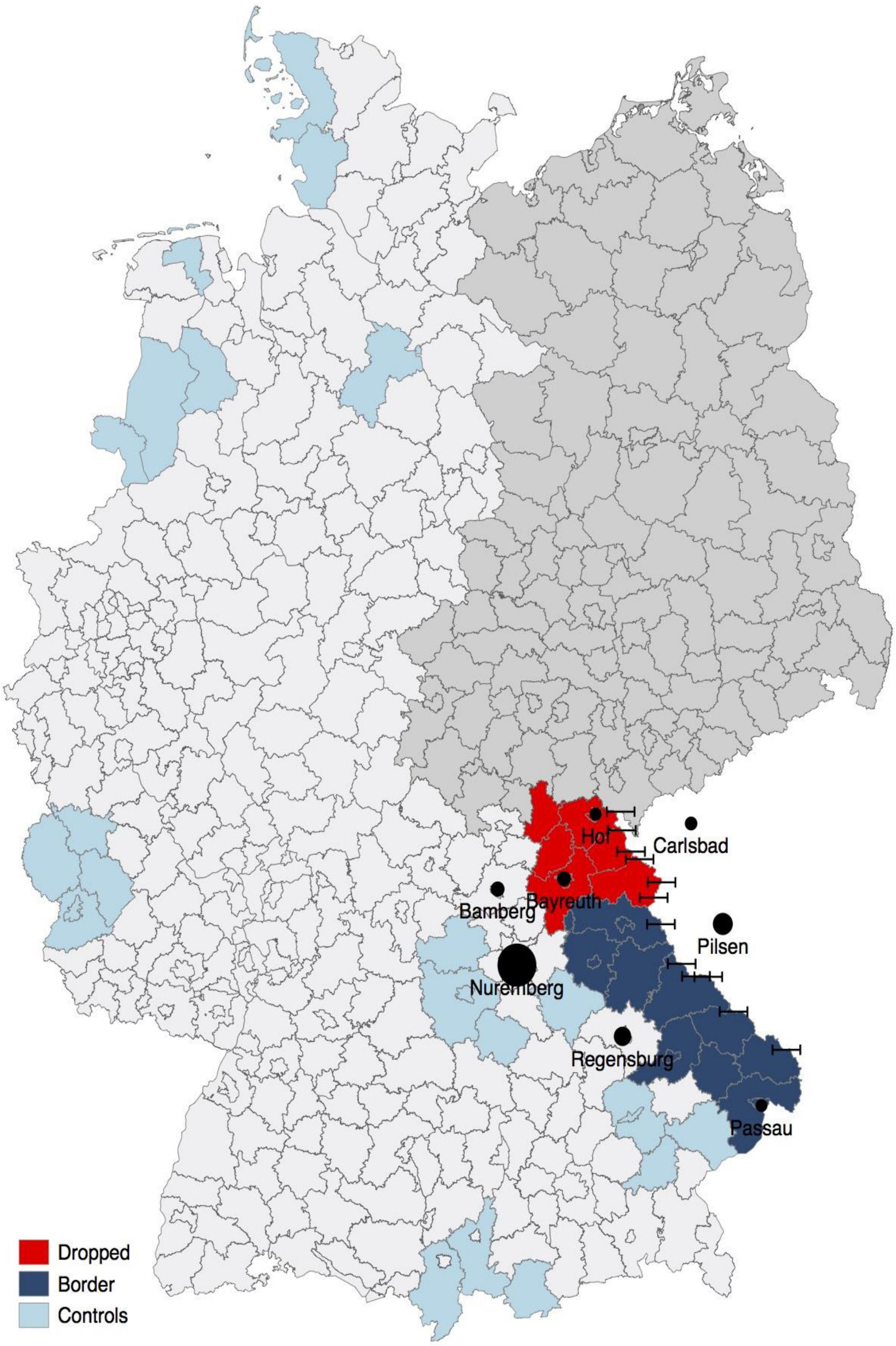
Figure 1: Border Region



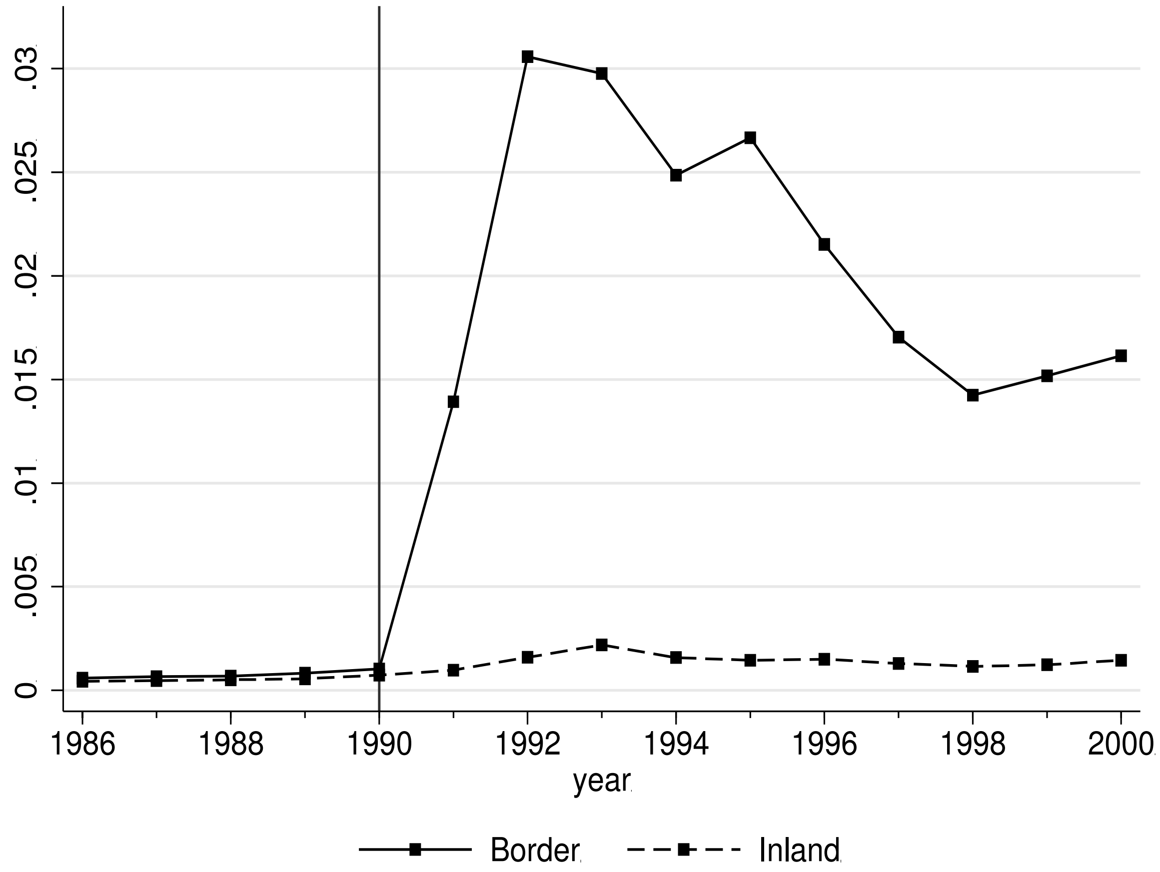
Note: The map shows districts eligible under the commuting policy (dark blue and red), matched inland control districts (medium blue), and other districts in West (light grey) and former East (darker grey) Germany. Eligible districts close to the inner German border (dark red) are dropped in the analysis. The map also shows crossings along and cities near the Czech-German border.

Table 1: Characteristics of treated, inland and matched control districts in 1989

	border	West Germany	control districts
skill			
low (no post-secondary education)	0.274	0.229	0.244
medium (apprenticeship or equivalent)	0.695	0.703	0.723
high (university or college)	0.030	0.069	0.034
age			
below 30	0.434	0.351	0.420
30 to 49	0.410	0.454	0.412
50 and above	0.157	0.195	0.168
female	0.411	0.401	0.414
foreign	0.025	0.081	0.035
mean log wages (censored)	3.881	4.055	3.879
share censored	0.023	0.048	0.027
# districts	13	327	24
# workers	335,042	21,173,830	726,536

Note: The table compares average characteristics (weighted by employment level) of workers in eligible districts in the border region, in all other West-German districts and in matched control districts (see Figure 1) in 1989, one year prior to the immigration-induced labor supply shock. The wage variable refers to the average wage earned per day of the employment relationship and right-censored at the social security limit.
Data Source: German Social Security Data, 1989.

Figure 2: Employment Shares of Czech nationals: Border vs Inland



Note: The figure plots the share of Czech workers in local employment in the border region and in matched control districts (see Figure 1) before and after the commuting policy came into effect in 1991.
Data Source: German Social Security Records, eligible border region and matched control districts, 1986 to 2000.

Table 2: Characteristics of Czech and Non-Czech Nationals in the Border Region

Panel A: Non-Czechs vs Czechs	Non-Czech (1989)	Czech (1992)
skill distribution		
unskilled (no post-secondary education)	0.276	0.505
skilled (apprenticeship or equivalent, college, or university)	0.724	0.495
age distribution		
below 30	0.435	0.344
30-49	0.408	0.619
50 and above	0.157	0.037
age distribution: unskilled		
below 30	0.500	0.370
30-49	0.290	0.593
50 and above	0.209	0.037
age distribution: skilled		
below 30	0.410	0.317
30-49	0.453	0.646
50 and above	0.137	0.037
share female	0.411	0.163
industries		
public sector	0.171	0.021
pit and quarry	0.027	0.048
wood processing	0.032	0.074
construction	0.099	0.249
hotels and restaurants	0.031	0.092
# workers	332,785	9,996

Panel B: Relative Wage Gap Czechs vs Non-Czechs (1992)	*Coeff.*	*S.E.*
(i) municipality fixed effects	-0.354	(0.003)
(ii) municipality and occupation fixed effects	-0.304	(0.003)
(iii) occupation X firm fixed effects	-0.224	(0.003)
# workers	267,756	

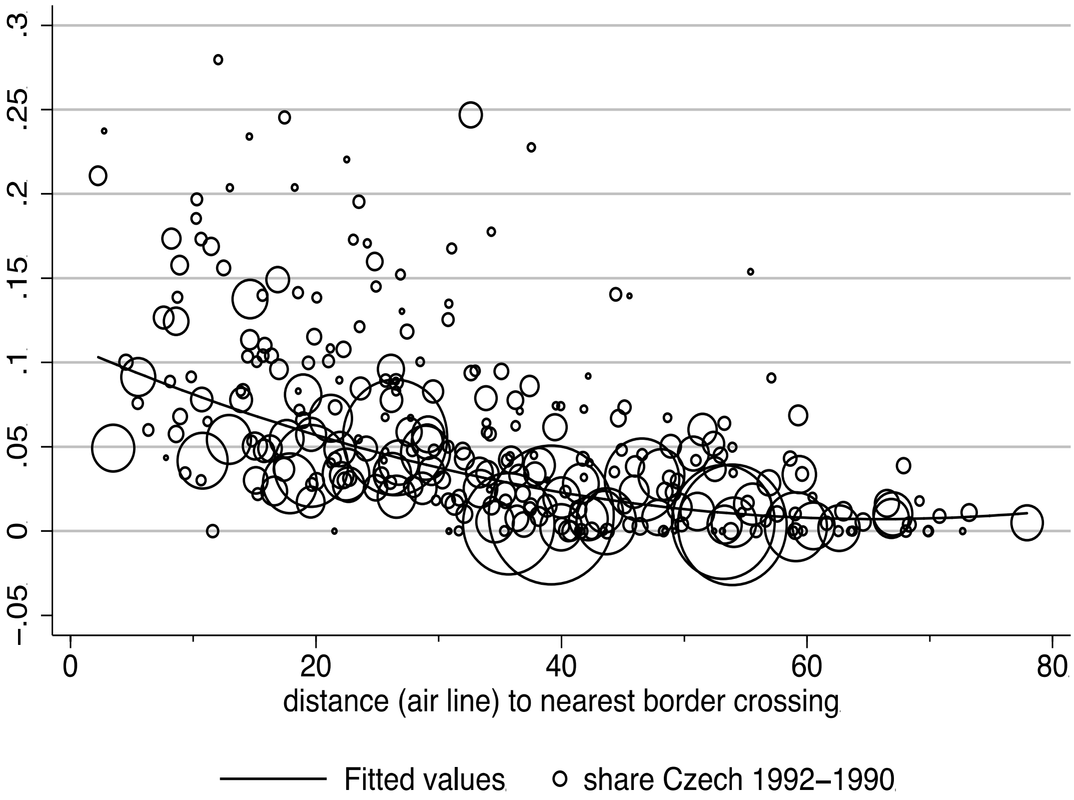
Note: Panel A compares the characteristics of Czech commuters (in 1992) against the pre-existing, non-Czech workforce (in 1989). Panel B reports the log-wage gap between Czech and Non-Czech workers in the border region in 1992. The wage variable refers to the average wage earned per day of the employment relationship and right-censored at the social security limit. Following Dustmann et al. (2009), we impute censored wages under the assumption that the error term is normally distributed while allowing for a different residual variance by gender as well as by district. All regressions control for age, age squared, and sex. Row (i) additionally controls for municipality fixed effects. Row (ii) further adds 3-digit occupation fixed effects. Row (iii) controls for 3-digit occupation X firm fixed effects. Robust standard errors in parentheses.
Data Source: German Social Security Records, eligible border region, 1989 and 1992.

Table 3: Placebo Regressions of Pre-Shock Employment and Wage Growth on Distance to Border

	employment growth			log wage growth		
	(1)	(2)	(3)	(4)	(5)	(6)
	all	unskilled	skilled	all	unskilled	skilled
Panel A: all						
	-0.023	-0.018	-0.081	0.005	-0.001	0.005
	(0.048)	(0.064)	(0.047)	(0.007)	(0.012)	(0.005)
Panel B: by age						
below 30	-0.013	0.016	-0.064	-0.001	-0.012	0.000
	(0.063)	(0.078)	(0.062)	(0.010)	(0.020)	(0.010)
30-49	-0.058	-0.074	-0.121*	0.010	0.004	0.013
	(0.049)	(0.076)	(0.049)	(0.008)	(0.017)	(0.007)
50 and above	0.031	0.011	-0.024	0.015	0.012	0.017
	(0.052)	(0.064)	(0.064)	(0.009)	(0.013)	(0.010)
# municipalities	291	291	291	291	291	291

Note: The table reports the slope coefficient from a regression of native employment and wage growth between 1987 and 1989 in the municipality on the municipality's airline distance to border (measured in km/100), separately for skill and age. The sample is restricted to the border region. Standard errors are clustered at the district level.
Data Source: German Social Security Records, border region, 1987 to 1989.

Figure 3: Spatial Distribution of Czech Commuters in Border Region



Note: The figure plots, for each municipality within the border region, the increase in the number of Czech workers as a share of employment in 1990 against the airline distance of the centroid of the municipality to the closest border crossing. The size of each circle is proportional to employment in 1990. Fitted values are from a regression on distance and distance squared.

Data Source: German Social Security Records, border region, 1990 and 1992.

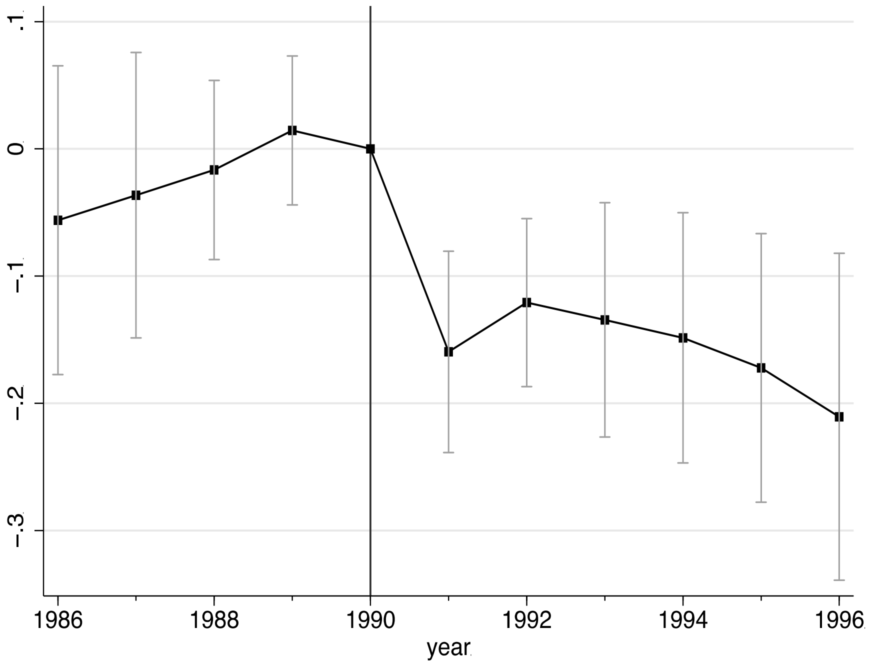
Table 4: First Stage: The Inflow of Czech Commuters and Distance to Border

	border region only	including matched control districts
distance(x100)	-0.338	-0.338
	(0.095)	(0.092)
distance(x100) squared	0.268	0.268
	(0.113)	(0.110)
constant (border region)	0.115	0.114
	(0.017)	(0.016)
constant (inland)		0.0011
		(0.0003)
# municipalites	291	1550
R-sq	0.387	0.544
F	42.58	52.70

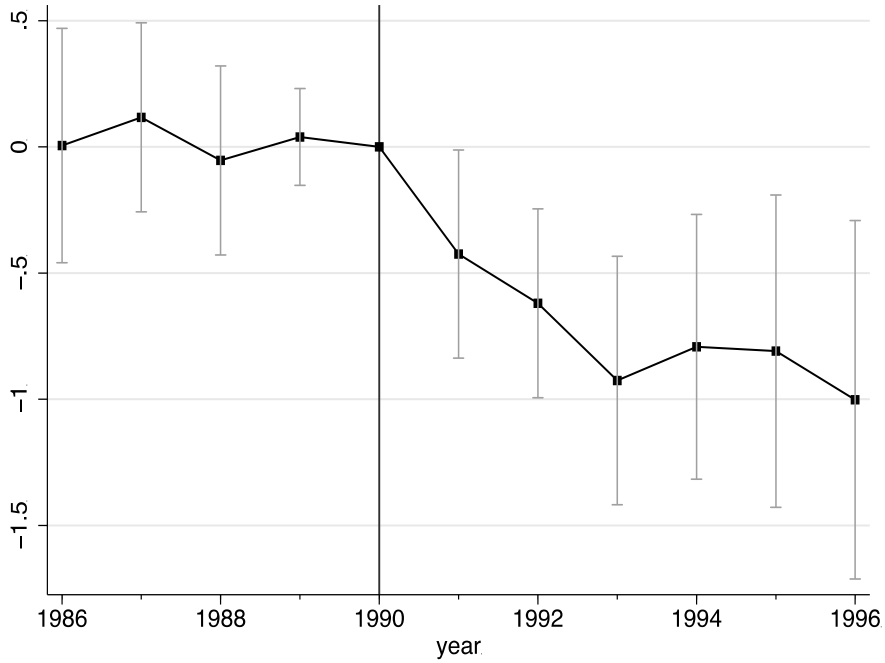
Note: The table reports the coefficients from the first stage regression of the inflow of Czech workers into the municipality, measured as the increase in the number of Czech workers between 1990 and 1992 as a share of local employment in 1990, on airline distance and distance squared to the next border crossing. Regressions are estimated at the municipality level, weighted by local employment in 1990. In the first column, the sample is restricted to the border region. The second column additionally includes matched control districts, and distance and distance squared is interacted with an indicator variable equal to 1 if the municipality is part of the border region. Standard errors are clustered on the district level.
Data Source: German Social Security Records, border region and matched control districts, 1990 and 1992.

Figure 4: Aggregate Wage and Employment Effects

Panel A: Wage Effects



Panel B: Employment Effects



Note: The figures are based on equations (6) and (7), where we regress at the municipality level the change in native log-wages or the percentage change in employment between two consecutive years on the inflow of Czech workers between 1990 and 1992 (except for the entry in year 1991 when we use the inflow of Czech workers between 1990 and 1991), instrumented by the municipality's distance to the border (and its square). We then plot the cumulative effects, starting in 1990, by adding up estimated coefficients backwards and forwards. While the first stage regression is weighted by native employment in the municipality in 1990, the second stage regressions are weighted by native employment in the respective base year. The 95% confidence intervals are computed using the wild bootstrap method, using 500 replications, allowing for clustering on the district level.

Data Source: German Social Security Records, border region and matched control districts, 1986 to 1996.

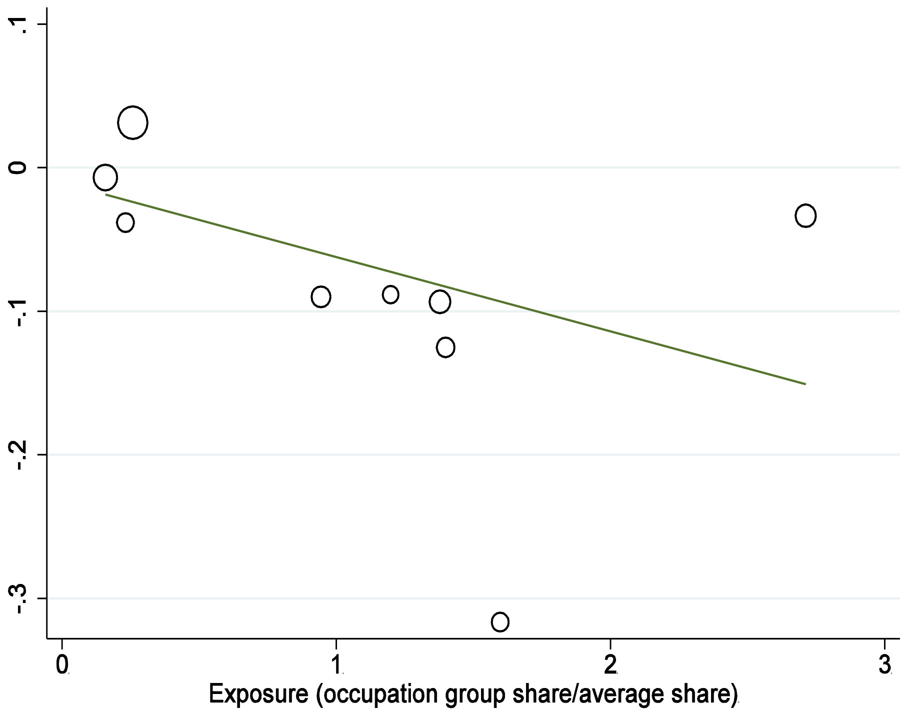
Table 5: Wage and Employment Baseline Estimates by Skill, 1990-1993

		wages	employment
Panel A: all			
(i)	2SLS	-0.134	-0.926
		(0.047)	(0.251)
(ii)	OLS	-0.058	-0.263
		(0.038)	(0.184)
Panel B: unskilled			
(i)	2SLS	-0.202	-1.371
		(0.048)	(0.395)
(ii)	OLS	-0.094	-0.789
		(0.041)	(0.215)
Panel C: skilled			
(i)	2SLS	-0.106	-0.501
		(0.051)	(0.214)
(ii)	OLS	-0.054	0.049
		(0.025)	(0.196)
	# municipalities	1,550	1,550

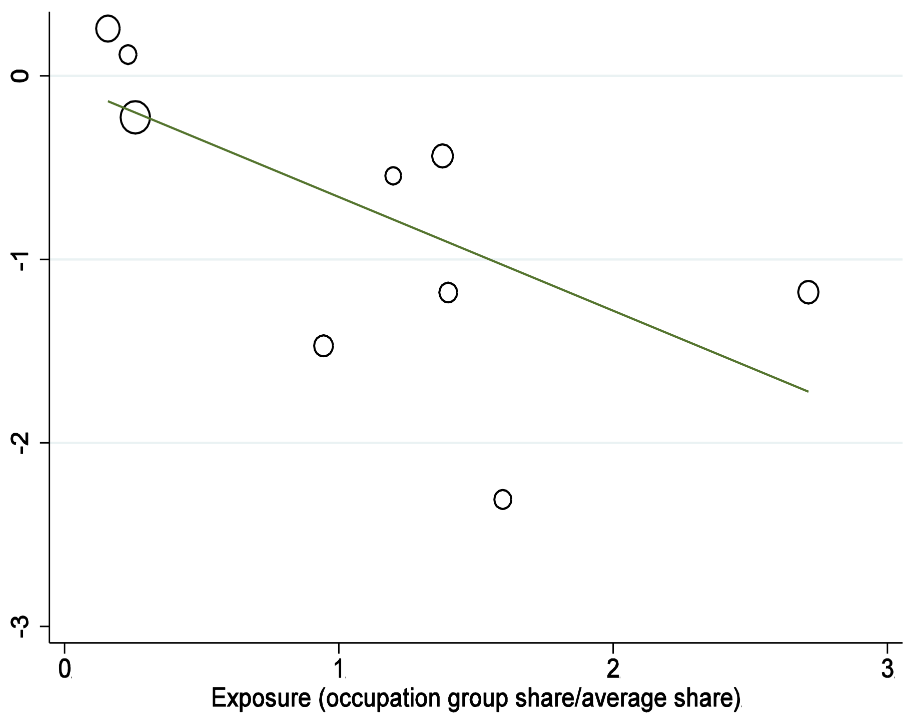
Note: The table reports 2SLS (rows (i)) and OLS (rows (ii)) estimates for the impact of the inflow of Czech commuters into the municipality, measured as the increase in the number of Czech workers between 1992 and 1990 as of employment in 1990, on native local wage and employment growth in the aggregate (Panel A) and for unskilled and skilled natives (Panels B and C). In rows (i), the inflow of Czech workers is instrumented with a quadratic in the municipality's airline distance to the nearest border crossing. Regressions are estimated at the yearly level, across up to N=1,550 municipalities, and coefficients are added up to obtain cumulative effects. To make sure that the wage effects are not underestimated because of worker selection, the yearly wage growth regressions are restricted to workers who remain employed in the district between two consecutive years. While the first stage regression is weighted by total native employment in the municipality in 1990, the second stage regression is weighted by group-specific native employment in the respective base year. Standard errors are bootstrapped, using 500 replications, allowing for clustering on the district level.
Data Source: German Social Security Records, border region and matched control districts, 1990 to 1993.

Figure 5: Wage and Employment Effects by Occupation Groups

Panel A: Wages



Panel B: Employment



The figure plots 2SLS estimates for the impact of the inflow of Czechs into the municipality on local native wage and employment growth in 9 one-digit occupations against the exposure to Czechs in these occupations, measured as the occupation-specific share of Czech workers relative to the mean share. Results refer to our baseline specification as in Table 4, rows (i).
Data Source: German Social Security Records, border region and matched inland control districts, 1990 to 1993.

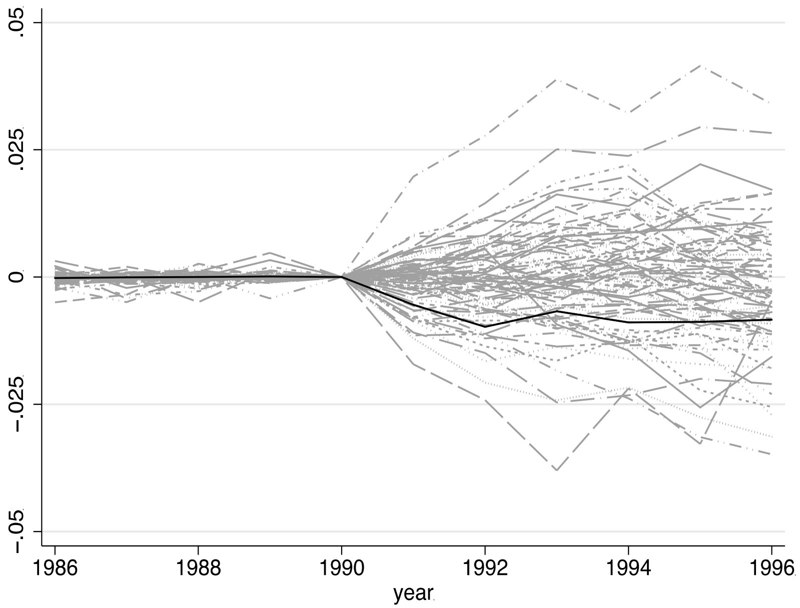
Table 6: Robustness Checks

	(1) baseline	(2) trend-adjusted	(3) border region only	(4) inner border region	(5) difference, clustered s.e.	(6) difference, s.e.
Panel A: Wage Effects						
all	-0.134	-0.209	-0.134	-0.142	0.002	0.002
	(0.047)	(0.056)	(0.096)	(0.050)	(0.066)	(0.053)
unskilled	-0.202	-0.282	-0.303	-0.205	-0.057	-0.057
	(0.048)	(0.068)	(0.105)	(0.051)	(0.106)	(0.080)
skilled	-0.106	-0.190	-0.093	-0.114	-0.052	-0.052
	(0.051)	(0.060)	(0.098)	(0.054)	(0.057)	(0.050)
Panel B: Employment Effects						
all	-0.926	-0.927	-0.952	-0.897	-0.930	-0.930
	(0.251)	(0.311)	(0.456)	(0.275)	(0.243)	(0.243)
unskilled	-1.371	-1.417	-1.036	-1.368	-1.203	-1.203
	(0.395)	(0.411)	(0.522)	(0.382)	(0.328)	(0.271)
skilled	-0.501	-0.866	-0.586	-0.507	-0.522	-0.522
	(0.214)	(0.313)	(0.450)	(0.236)	(0.214)	(0.230)
# municipalities	1550	1550	291	1405	1550	1550

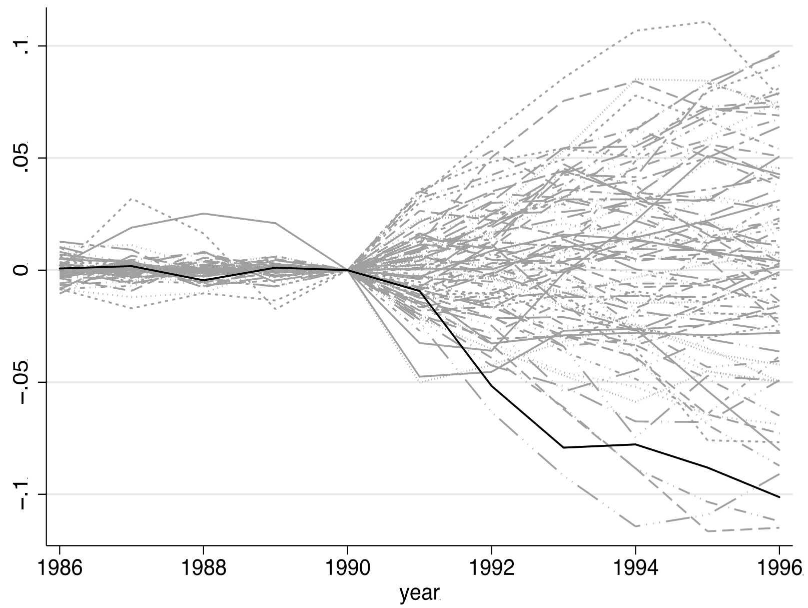
Note: The table presents coefficient estimates from various robustness tests. Column (1) reports our baseline estimates (see Table 4, rows (i)). Column (2) allows for linear municipality-specific time trends in pooled regressions, in which the pre-treatment observations in 1987-1989 identify municipality-specific differences in trend. Column (3) drops matched control districts from the sample and uses variation in the inflow of Czechs across municipalities within the border region only. Column (4) compares the highly affected Eastern (“inner”) part of the border region to unaffected matched control districts. In columns (5) and (6), we report estimates for which we take long differences (between 1990 and 1993) and average log wages over all workers who are in employment in either of the two years, rather than over workers who remain employed in two consecutive years, as in our baseline specification in column (1). In columns (1) to (5), standard errors are bootstrapped, using 500 replications, allowing for clustering on the district level. Column (6) displays instead standard error estimates based upon spatial HAC technique of Conley (1999), using a uniform kernel and bandwidth of 100 kilometres.
Data Source: German Social Security Records, border region and matched control districts, 1987 to 1993.

Figure 6: Synthetic Control Method, Employment and Wage Effects

Panel A: Wages



Panel B: Employment



Note: The figures display the wage (Panel A) and employment (Panel B) gap between the highly exposed inner border region and its synthetic control (bold line) and the respective placebo gaps from 85 control districts, following Abadie et al. (2010); see Appendix C for details. To make sure that the wage effects are not underestimated because of worker selection, wage growth is computed for all workers who remain employed in the district between two consecutive years, as in our baseline specification. To compare results from the synthetic control approach with our baseline estimates, the figures also display the scaled-up gaps, where we divide the employment and wage gaps between the inner border region and its synthetic control in 1993 by the average increase in the number of Czech workers in the inner border region as of its employment in 1990 (0.058).
Data Source: German Social Security Data, inner border region and inland districts, 1986 to 1996.

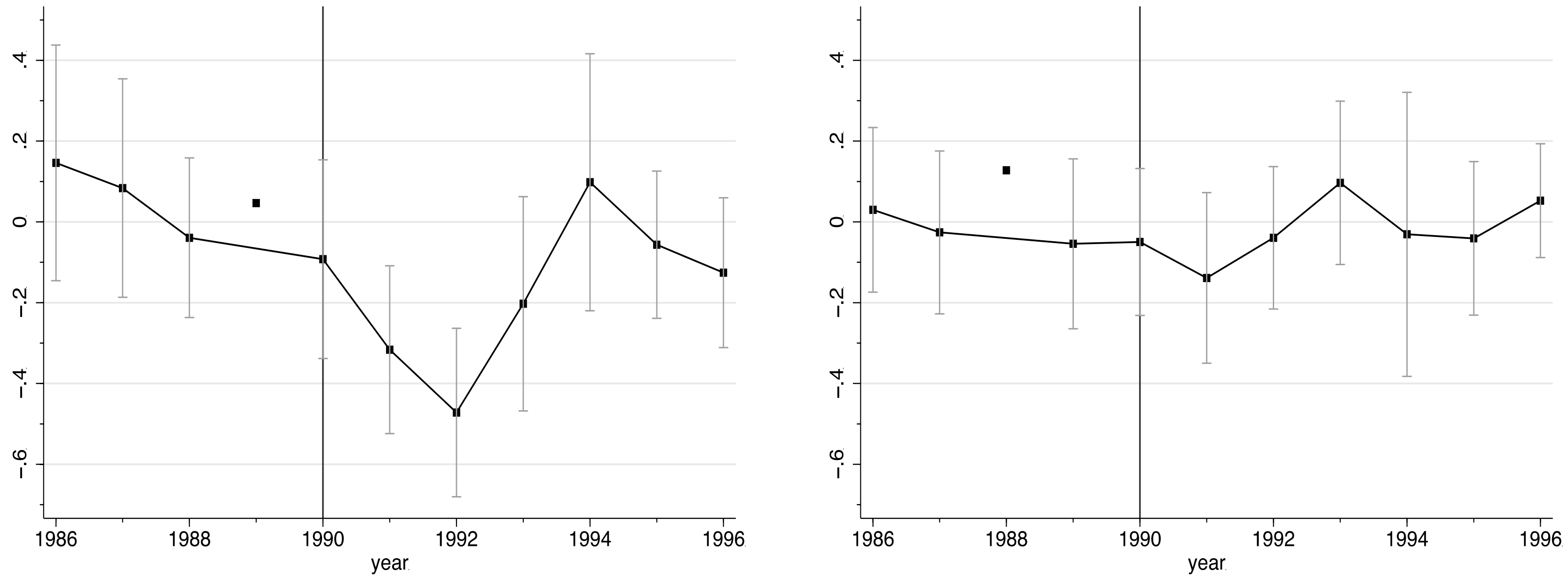
Table 7: Employment and Wage Effects by Skill and Age Groups

	(1)	(2)	(3)	(4)
	unadjusted		trend-adjusted	
	wages	employment	wages	employment
Panel A: all				
below 30	-0.316	-0.832	-0.305	-0.604
(share Czechs: 0.031)	(0.086)	(0.317)	(0.079)	(0.373)
30 to 49	-0.100	-0.534	-0.147	-0.964
(share Czechs: 0.040)	(0.050)	(0.238)	(0.058)	(0.338)
50 and above	-0.068	-1.945	-0.172	-1.428
(share Czechs: 0.007)	(0.046)	(0.340)	(0.055)	(0.394)
Panel B: unskilled				
below 30	-0.558	-2.262	-0.441	-1.601
(share Czechs: 0.112)	(0.107)	(0.585)	(0.103)	(0.549)
30 to 49	-0.179	-0.704	-0.237	-1.428
(share Czechs: 0.107)	(0.064)	(0.428)	(0.078)	(0.501)
50 and above	-0.097	-1.364	-0.194	-1.324
(share Czechs: 0.011)	(0.053)	(0.342)	(0.080)	(0.470)
Panel C: skilled				
below 30	-0.276	-0.283	-0.281	-0.457
(share Czechs: 0.017)	(0.092)	(0.319)	(0.081)	(0.378)
30 to 49	-0.090	-0.191	-0.142	-1.012
(share Czechs: 0.025)	(0.058)	(0.197)	(0.063)	(0.329)
50 and above	-0.066	-1.636	-0.158	-1.337
(share Czechs: 0.005)	(0.053)	(0.275)	(0.061)	(0.383)
# municipalities	1,550	1,550	1,550	1,550

Note: The table reports 2SLS estimates for the cumulative impact, between 1990 and 1993, of the inflow of Czech commuters between 1990 and 1992 on local wage and employment growth of natives by age (Panel A) and by age and skill (Panels B and C). Columns (1) and (2) report unadjusted estimates from regressions estimated at the yearly level, where coefficients are added up to obtain cumulative effects. Columns (3) and (4) report trend-adjusted estimates, obtained from a pooled regressions over the years 1987 to 1993, in which pre-treatment observations in 1987-1989 identify differences in the linear municipality-specific time trend. While the first stage regression is weighted by total native employment in the municipality in 1990, the second stage regression is weighted by group-specific native employment in the respective base year. Standard errors are bootstrapped, using 500 replications, allowing for clustering at the district level.
Data Source: German Social Security Records, border region and matched inland control districts, 1987 to 1993.

Figure 7: Yearly native inflow and outflow effects

Panel A: Inflow Effects **Panel B: Outflow Effects**



Note: The figures plot coefficient estimates from the 2SLS regressions of yearly inflow or outflow rates of natives on the inflow of Czech workers in the municipality between 1990 and 1992. While the first stage regression is weighted by total native employment in the municipality in 1990, the second stage regression is weighted by native employment in the respective base year. The 95% confidence interval is based on bootstrapped standard errors which use 500 replications and allow for clustering on the municipality level. The coefficient estimate for outflows in 1989 and inflows in 1988 represent outliers (see details in the text) and are plotted, but not connected to the coefficient estimates.

Data Source: German Social Security Records, border region and matched inland control districts, 1986 to 1996.

Table 8: Margins of Adjustment: Inflows vs Outflows and Geographical vs Non-employment Movements

		inflows vs outflows		geographical vs non-employment movements			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	total employment	inflows	outflows	non-employment	geographic., direct	geographic., incl. through non-employment	population
Panel A: all							
share of baseline employment		*0.18*	*0.16*	*0.11*	*0.06*	*0.08*	
	-0.989	-0.878	0.111	-0.821	-0.168	-0.287	-0.299
	(0.318)	(0.258)	(0.152)	(0.229)	(0.169)	(0.182)	(0.059)
Panel B: by skill							
unskilled							
share of baseline employment		*0.15*	*0.17*	*0.12*	*0.04*	*0.05*	
	-1.256	-1.385	-0.129	-1.289	0.033	-0.103	
	(0.534)	(0.391)	(0.210)	(0.422)	(0.201)	(0.242)	
skilled							
share of baseline employment		*0.19*	*0.16*	*0.11*	*0.06*	*0.08*	
	-0.875	-0.761	0.115	-0.658	-0.218	-0.326	
	(0.290)	(0.250)	(0.144)	(0.217)	(0.169)	(0.178)	
Panel C: by age							
below 30							
share of baseline employment				*0.17*	*0.09*	*0.12*	
	-0.555	-0.594	-0.039	-0.560	-0.147	-0.254	
	(0.416)	(0.369)	(0.163)	(0.328)	(0.174)	(0.208)	
30 to 49							
share of baseline employment				*0.07*	*0.05*	*0.06*	
	-1.180	-1.330	-0.150	-0.857	-0.257	-0.381	
	(0.382)	(0.281)	(0.177)	(0.187)	(0.220)	(0.238)	
50 and above							
share of baseline employment				*0.10*	*0.02*	*0.03*	
	-1.349	-0.974	0.375	-1.203	-0.050	-0.061	
	(0.441)	(0.352)	(0.303)	(0.317)	(0.164)	(0.185)	
# municipalities	1,550	1,550	1,550	1,550	1,550	1,550	1,550

Note: The table first breaks down the overall cumulative (from 1990 to 1993) local employment effect of natives (reported in column (1)) into inflows from either non-employment or employment in other areas (column (2)) and outflows into either non-employment or employment in other areas (column (3)), in the aggregate (Panel A) and separately by skill (Panel B) and age (Panel C). Columns (4) and (5) split up the overall employment effect into direct non-employment and direct geographical movements (i.e., transitions from employment in one municipality in one period and in another in the next). Column (6) allows for geographical movements through non-employment, and re-categorizes inflows from and outflows to non-employment as geographic movement if within three years the worker is observed working in a different municipality. Column (7) displays the 2SLS estimate for impact of the inflow of Czechs between 1990 and 1992 into the municipality on population growth (data obtained from the German Federal Statistical Office) in the municipality between 1990 and 1993. Estimates are trend-corrected using years 1989 and 1990 to account for pre-existing linear municipality-specific time trends. The table also reports shares of inflows and outflows, as well as shares of non-employment and geographical movements, averaged over the pre-policy years 1985 to 1989 in italics. Note that in Panel C, figures do not exactly add up to the total employment effect as there are inflows and outflows across age groups. Standard errors are bootstrapped using 500 replications, and allow for clustering on the district level.
Data Source: German Social Security Records, border region and matched inland control districts, 1989 to 1993.

Table 9: Inflows and Outflows into and from Other areas and Non-Employment

		inflows		outflows	
	(i)	(ii)	(iii)	(iv)	(v)
	total employment	inflows into non-employment (direct)	inflows into employment in other areas (direct)	outflows into non-employment (direct)	outflows into employment in other areas (direct)
Panel A: all					
	-0.989	-0.645	-0.233	0.176	-0.066
	(0.318)	(0.207)	(0.173)	(0.081)	(0.123)
Panel B: by skill					
unskilled	-1.256	-1.154	-0.231	0.135	-0.264
	(0.534)	(0.326)	(0.187)	(0.176)	(0.137)
skilled	-0.875	-0.498	-0.263	0.159	-0.045
	(0.290)	(0.196)	(0.182)	(0.068)	(0.129)
Panel C: by age					
below 30	-0.555	-0.315	-0.280	0.245	-0.133
	(0.416)	(0.347)	(0.171)	(0.108)	(0.144)
30 to 49	-1.180	-0.807	-0.275	0.050	-0.018
	(0.382)	(0.176)	(0.190)	(0.057)	(0.159)
50 and above	-1.349	-0.726	-0.051	0.477	-0.101
	(0.441)	(0.215)	(0.181)	(0.244)	(0.152)
# municipalities	1,550	1,550	1,550	1,550	1,550

Note: The table breaks up the overall cumulative employment effect between 1990 and 1993 into direct inflows from (column (1)) and outflows to (column (4)) non-employment, and direct inflows to (column (3)) and outflows from (column (5)) employment in one municipality to another. Estimates are trend-corrected using years 1989 and 1990 to account for pre-existing linear municipality-specific time trends. Standard errors are bootstrapped using 500 replications, and allow for clustering on the district level.
Data Source: German Social Security Records, border region and matched inland control districts, 1989 to 1993.

Labor Supply Shocks, Native Wages, and the Adjustment of Local Employment

Online Appendix

Christian Dustmann, Uta Schönberg, and Jan Stuhler

Appendix A: Model

A.I: Derivation of the Firm's Demand Curve

From $K = h(r, \boldsymbol{r}')$, and $\frac{1}{\lambda} = \frac{\partial h}{\partial r}\frac{r}{h}$, we obtain

$$\mathrm{d}logr = \lambda\, dlogK. \tag{A.1}$$

Totally differentiating Equation (1b) and plugging in Expression (A.1) for $\mathrm{d}logr$ yields

$$dlogK = -\frac{\alpha-1}{1-\alpha+\lambda} dlogL. \tag{A.2}$$

Totally differentiating Equation (1a) and substituting for $log\ K$ using (A.2) then yields

$$dlogw_g = \varphi dlogL + (\beta - 1)\left(dlogL_g - dlogL\right), \tag{A.3}$$

where

$$\varphi = -\frac{\alpha\lambda}{1-\alpha+\lambda} \tag{A.4}$$

is the slope of the aggregate labor demand curve.

Starting from $L_g = L_g^I + L_g^N$ and assuming that there are no immigrants at baseline, we obtain $dlogL_g = \frac{dL_g^I}{L} + dlogL_g^N$. Letting $\pi_g^N = \frac{L_g^N}{L_U^N + L_S^N}$ and $\pi_g^I = \frac{L_g^I}{L_U^I + L_S^I}$ to denote the employment shares of natives and immigrants in skill group g (in head count), the expression above can be rewritten as

$$dlogL_g = \frac{\pi_g^I}{\pi_g^N} dI + dlogL_g^N. \tag{A.5}$$

Totally differentiating $L = \left[\theta_U L_U^\beta + \theta_S L_S^\beta\right]^{\frac{1}{\beta}}$ results in $dlogL = s_U dlogL_U + s_S dlogL_S$, where $s_g = \frac{\theta_g L_g^\beta}{\left[\theta_U L_U^\beta + \theta_S L_S^\beta\right]}$ are the contribution of labor type g to the total labor aggregate. Substituting for $dlogL_U$ and $dlogL_S$ using expression (A.5) results in

$$dlogL = \Pi dI + s_U dlogL_U^N + s_S dlogL_S^N, \tag{A.6}$$

where $\Pi = s_U \frac{\pi_U^I}{\pi_U^N} + s_S \frac{\pi_S^I}{\pi_S^N}$ is the weighted average of the relative density of immigrants across skill groups. Plugging in expressions (A.5) and (A.6) for $dlogL$ and $dlogL_g$ in expression (A.3) and rearranging results in

$$dlogL_g^N = \frac{dlogw_g - \big(\varphi - (\beta-1)\big)s_{g'} dlogL_{g'}^N - \left(\big(\varphi - (\beta-1)\big)\Pi + (\beta-1)\frac{\pi_g^I}{\pi_g^N}\right)dI}{\varphi s_g + (\beta-1)(1-s_g)} \tag{A.7}$$

and a corresponding equation for the other skill group g'. Plugging these equations into each other leads to equation (2) in the text.

A.II: Derivation of the Equilibrium Wage and Employment Responses under Flexible Wages

The equilibrium wage and employment responses are determined by the two skill-specific labor demand curves (see equation A.3)

$$dlogw_U = \varphi dlogL + (\beta - 1)(dlogL_U - dlogL) \qquad \text{(A.8)}$$

$$dlogw_S = \varphi dlogL + (\beta - 1)(dlogL_S - dlogL) \qquad \text{(A.9)}$$

and the two skill-specific supply curves

$$d\mathrm{log}L_U^N = \eta_U d\mathrm{log}w_U \qquad \text{(A.10)}$$

$$d\mathrm{log}L_S^N = \eta_S d\mathrm{log}w_{S,} \qquad \text{(A.11)}$$

where $dlogL$ is given by Equation (A.6). By plugging (A.10) and (A.11) into (A.8) and (A.9), we obtain

$$dlogw_U = \varphi\big((s_U\eta_U dlogw_U + s_S\eta_S dlogw_S) + \Pi dI\big)$$

$$+(\beta - 1)\left(\eta_U dlogw_U - (s_U\eta_U dlogw_U + s_S\eta_S dlogw_S) + \left(\frac{\pi_U^I}{\pi_U^N} - \Pi\right) dI\right) \qquad \text{(A.12)}$$

$$dlogw_S = \varphi\big((s_U\eta_U dlogw_U + s_S\eta_S dlogw_S) + \Pi dI\big)$$

$$+(\beta - 1)\left(\eta_S dlogw_S - (s_U\eta_U dlogw_U + s_S\eta_S dlogw_S) + \left(\frac{\pi_S^I}{\pi_S^N} - \Pi\right) dI\right). \qquad \text{(A.13)}$$

Solving (A.12) and (A.13) for $dlogw_U$ and $dlogw_S$, respectively, gives

$$dlogw_U = \frac{\big(\varphi-(\beta-1)\big)s_S\eta_S dlogw_S + \varphi\Pi dI + (\beta-1)\left(\frac{\pi_U^I}{\pi_U^N}-\Pi\right)dI}{1-\varphi s_U\eta_U-(\beta-1)s_S\eta_U} \tag{A.14}$$

$$dlogw_S = \frac{\big(\varphi-(\beta-1)\big)s_U\eta_U dlogw_U + \varphi\Pi dI + (\beta-1)\left(\frac{\pi_S^I}{\pi_S^N}-\Pi\right)dI}{1-\varphi s_S\eta_S-(\beta-1)s_U\eta_S}. \tag{A.15}$$

Plugging (A.15) into (A.14) and placing all terms over a common denominator then yields

$$dlogw_U = \frac{\big(\varphi-(\beta-1)\big)^2 s_U s_S\eta_U\eta_S dlogw_U + \big(\varphi-(\beta-1)\big)s_S\eta_S\left(\varphi\Pi+(\beta-1)\left(\frac{\pi_S^I}{\pi_S^N}-\Pi\right)\right)}{(1-\varphi s_U\eta_U-(\beta-1)s_S\eta_U)(1-\varphi s_S\eta_S-(\beta-1)s_U\eta_S)}dI$$
$$+\frac{(1-\varphi s_S\sigma_S-(\beta-1)s_U\eta_S)\varphi\Pi+(1-\varphi s_S\eta_S-(\beta-1)s_U\eta_S)(\beta-1)\left(\frac{\pi_U^I}{\pi_U^N}-\Pi\right)}{(1-\varphi s_U\eta_U-(\beta-1)s_S\eta_U)(1-\varphi s_S\eta_S-(\beta-1)s_U\eta_S)}dI.$$

and solving for $dlogw_U$ gives

$$dlogw_U = \frac{\big(\varphi-(\beta-1)\big)s_S\eta_S\left(\varphi\Pi+(\beta-1)\left(\frac{\pi_S^I}{\pi_S^N}-\Pi\right)\right)}{(1-\varphi s_U\eta_U-(\beta-1)s_S\eta_U)(1-\varphi s_S\eta_S-(\beta-1)s_U\eta_S)-\big(\varphi-(\beta-1)\big)^2 s_U s_S\eta_U\eta_S}dI$$
$$+\frac{(1-\varphi s_S\eta_S-(\beta-1)s_U\eta_S)\varphi\Pi+(1-\varphi s_S\eta_S-(\beta-1)s_U\eta_S)(\beta-1)\left(\frac{\pi_U^I}{\pi_U^N}-\Pi\right)}{(1-\varphi s_U\eta_U-(\beta-1)s_S\eta_U)(1-\varphi s_S\eta_S-(\beta-1)s_U\eta_S)-\big(\varphi-(\beta-1)\big)^2 s_U s_S\eta_U\eta_S}dI.$$

Simplifying both the numerators and the denominator and using $\phi = \frac{\varphi}{\beta-1} - 1$ then leads to the equilibrium wage and employment responses given by equations (4) and (5) in the text.

A.III: Wage and Employment Responses for Three Skill Groups

In the case of two types of labor, perverse effects will only emerge when capital is not perfectly elastic (i.e., $\varphi < 0$) because given a perfectly elastic capital supply, the aggregate wage effect of a migration-induced supply shock is zero and the wage decreases for the skill group having a higher migrant share and increases for the other skill group regardless of the relative magnitude of the group-specific local labor supply elasticities. By extending the model to three types of labor we show that perverse effects are possible even when capital is fully elastic (i.e., $\varphi = 0$).

Assume that labor L is a CES aggregate of low (*L*), medium (*M*), and high (*H*) skilled labor, such that

$$L = \left[\theta_L L_L^{\beta} + \theta_M L_M^{\beta} + \theta_H L_H^{\beta}\right]^{\frac{1}{\beta}}. \tag{A.16}$$

As before, $dlogL_g = \frac{\pi_g^I}{\pi_g^N} dI + dlogL_g^N$ (see Equation A.5), while Equation (A.6) becomes

$$dlogL = \Pi dI + s_L dlogL_L^N + s_M dlogL_M^N + s_H dlogL_H^N, \tag{A.17}$$

with $s_g = \frac{\theta_g L_g^{\beta}}{\left[\theta_L L_L^{\beta} + \theta_M L_M^{\beta} + \theta_H L_H^{\beta}\right]}$. Since $\varphi = 0$, Equation (A.3) simplifies to

$$dlogw_g = (\beta - 1)\left(dlogL_g - dlogL\right). \tag{A.18}$$

Plugging in the expressions for $dlogL_g$ and $dlogL$, exploiting that with fully flexible wages $dlogL_L^N = \eta_L dlogw_L$, and solving for $dlogw_g$, we obtain the following for skill group $g = L$

$$dlogw_L = \frac{(\beta-1)\left(\left(s_M(1-(\beta-1)\eta_H)\left(\frac{\pi_L^I}{\pi_L^N}-\frac{\pi_M^I}{\pi_M^N}\right)+s_H(1-(\beta-1)\eta_M)\left(\frac{\pi_L^I}{\pi_L^N}-\frac{\pi_H^I}{\pi_H^N}\right)\right)dI\right.}{1-(\beta-1)\#1+(\beta-1)^2\#2}, \tag{A.19}$$

where

$$\#1 = \left((1-s_L)\eta_L + (1-s_M)\eta_M + (1-s_H)\eta_H\right)$$

$$\#2 = \left((1-s_L-s_M)\eta_L\eta_M + (1-s_M-s_H)\eta_M\eta_H + (1-s_L-s_H)\eta_L\eta_H\right).$$

The employment response follows from

$$dlogL_L^N = \eta_L dlogw_L.$$

The wage and employment responses of the other skill groups follow accordingly.

Perverse wage effects are thus possible. Supposing that migrant concentration is high in skill group *L*, medium in skill group *M*, and low in skill group *H*,

$$\frac{\pi_L^I}{\pi_L^N} > \frac{\pi_M^I}{\pi_M^N} > \frac{\pi_H^I}{\pi_H^N}$$

if the local labor supply elasticity of the medium skilled is large relative to that of the low skilled, the wages of the former can still decline more than the wages of the latter (i.e., $dlogw_M < dlogw_L$). It is, however, not possible that wages of the high skilled (which must increase if capital is fully flexible) decline relative to wages of the low skilled (which decline).

Appendix B: The Commuting Policy

After World War II, the border between West Germany and former Czechoslovakia became heavily fenced, allowing no movement of goods, capital, or people across. The first opening of the border occurred in November 3, 1989, when Czechoslovakia allowed East Germans—who had gathered in great numbers in West Germany's embassy in Prague—to travel to West Germany. Shortly after, in November 17, the border opened up also for Czechs and Slovakians to travel. Beginning on December 11, the Czechoslovak fortifications on the West German border were dismantled, and from July 1990, any visa requirements for cross-border travel were abolished.

Against this backdrop, various schemes for the legal employment of foreign nationals in Germany were extended or introduced with effect on January 1, 1991.[1] The overall policy comprised a locally constrained scheme that received little public attention, the *Grenzgängerregelung* commuting scheme, which granted foreign nationals from neighboring countries the right to work in dependent employment in German border regions without granting them residency rights. These workers were required to commute daily from their country of origin *or* to work for a maximum of two days per week. Since the movement of labor was in principal unrestricted within the European Economic Community (bilateral agreements already covered tax and other issues on the western borders), this policy had consequences only on the eastern German borders with Poland and Czechoslovakia (from 1993, the Czech Republic).[2] The intended implementation of the policy along the Czech-German border was first reported in September 1990, only shortly before the scheme came into effect.[3] The initial provision specified only 18 districts, but a revision lists three additional districts (*Straubing*, *Deggendorf*, and *Straubing-Bogen*), leading us to consider all 21 districts as treated units, although our results are robust to exclusion of the additions. As explained in the text, we exclude 8 districts located within about 80 kilometers of the former border between East- and West Germany because external data on regional commuting flows from the late 1990s show that areas directly adjacent to that border received a high share of commuters.[4]

[1] See "Anwerbestoppausnahme-Verordnung", Bundesgesetzblatt, Jahrgang 1990, Teil I.
[2] A summary of the existing commuting schemes within the European Community is given in IAB (1993).
[3] See *Süddeutsche Zeitung* (1./2.9.1990). Implementation of the scheme along the Czech-German border occurred shortly before its general introduction on January 1, 1991 (see Hönekopp, 1991). A detailed examination of daily employment records confirms that the inflow of Czech nationals had already begun by September 1990.
[4] See, for example, Kropp (2010) and IAB (1992).

Czech commuters pay taxes in the Czech Republic—their country of residence—according to the Czech tax law. They pay social security contributions according to the German law. Yet, in case of a job loss, they are entitled to unemployment benefits only in the Czech Republic, but not in Germany. Because the commuting scheme had little effect on the West German labor market as a whole, it received little attention in national newspapers. In local newspapers, however, the increase in Czech commuters in the border region was perceived as having negative consequences for native workers.

Appendix C: Matching and the Synthetic Control Method

Matching of Control Units in Baseline Specification

To account for differences in district size, for each treated district, we match one or multiple control unit(s) until their employment levels sum to at least proportion x of employment in the treated unit. The choice of x is subject to a trade-off between bias and precision: choosing a higher value results in the matching of more but potentially less suitable control areas. For our baseline specification, we choose $x = 1.5$ but, to ensure that our findings are not driven by the particular choice of control units, we repeat the analysis for alternative sets of matched characteristics using other values of x. Doing so has little effect on our baseline findings.

Synthetic Control Approach

When applying the synthetic control method, we construct a comparison unit from all West-German districts located in rural regions or regions with intermediate agglomerations, and not neighboring the former border between East- and West Germany, according to the following procedure:

- First, we define the pre-intervention periods as between 1985 and 1990.
- Second, we define a vector $\boldsymbol{X_1}$ of preintervention characteristics and outcomes for the exposed region. In our case, $\boldsymbol{X_1}$ consists of the value of the outcome variable in each pre-intervention period and the average over the entire pre-intervention periods of (1) the employment growth among natives, (2) the wage growth among native incumbents, (3) the employment share of unskilled workers, (4) the share of foreign workers, and (5) the share of four age groups. We similarly define $\boldsymbol{X_0}$ as a matrix containing the same variables for the unaffected region.
- Third, we choose a weighting vector $\boldsymbol{W}^*$ to minimize the distance $\parallel \boldsymbol{X_1} - \boldsymbol{X_0 W} \parallel_{\boldsymbol{V}} = \sqrt{(\boldsymbol{X_1} - \boldsymbol{X_0 W})' \boldsymbol{V} (\boldsymbol{X_1} - \boldsymbol{X_0 W})}$, where $\boldsymbol{V}$ is selected from among the positive definite

and diagonal matrices such that the mean squared prediction error of the outcome variable is minimized for the pre-intervention periods.[5] The synthetic control method thus sets both a weight for each predictor (via $\boldsymbol{V}$) and a weight for each available control district (via $\boldsymbol{W}$).

Appendix D: Additional Results

D.1 Alternative First Stages

Our results are robust to alternative ways of estimating the first stage. In these robustness checks, we restrict the sample to the border region only and exclude the matched inland control districts (as all control districts are far from the border). For simplicity, we report estimates based on long differences instead of estimating annual regressions and adding up coefficients as in our baseline specification in the manuscript. These modifications have a negligible effect on the point estimates, but wage impacts are less precisely estimated.

First, we use different functional form assumptions for our distance measure, comparing our baseline estimate based on a quadratic of distance to border with estimates based on a third order polynomial and a spline function with 5, 10 or 20 knots in distance to border. The corresponding first-stage and 2SLS results for both employment and wages are reported in Panel A of Table O.1. 2SLS estimates hardly change when we use a polynomial of order 3 in distance to border to predict the inflow of Czech workers into the area (column (2)). 2SLS estimates are also very similar when a spline function with 5 knots is used to predict the inflow (column (3)). However, employment effects become less negative when more knots are included. We attribute this to an endogeneity problem: if we allow more flexibility in the first stage, it will increasingly capture that Czechs prefer to enter areas with strong labor demand; that is, labor demand in a band 20km from the border may have been somewhat higher than demand 25km away from the border.

Second, we use alternative distance measures (see Panel B of Table O.1). We computed driving distances in both km and time between all municipalities and the closest border crossing, using the Google Maps API. The correlation between airline distance (what we used

[5] See Appendix B in Abadie and Gardeazabal (2003). For implementation, we use the Synth software package provided by Abadie, Diamond, and Hainmueller at http://www.stanford.edu/~jhain/synthpage.html.

so far) and driving distances is very high (about 0.97) and accordingly, the 2SLS employment and wage estimates are similar to our baseline estimates (Column (2) in Panel B). The correlation between airline distances and driving times is slightly smaller (slightly less than 0.9), but once again the 2SLS wage coefficient differs only slightly from our baseline. As distances are computed based on the current (2015) road networks, it might be a worse proxy for actual driving time than simple airline distances. Moreover, the current road network may be endogenous to past differences in local labor conditions in general and the spatial distribution of Czech commuters in particular. For these reasons we prefer airline distance as our baseline measure.

D.2 Alternative Computations of Standard Errors

In Table O.2, we report additional estimates for the standard errors in the 2SLS wage and employment effects. As in Table O.1, the specification refers to the long difference (1990 to 1993) regression. The sample consists of the border region and matched inland control districts. In column (1), we report standard errors clustered on the district level, as in our baseline specification. In columns (2) to (4), we report standard errors based on Spatial Heteroscedasticity and Autocorrelation Consistent (SHAC) variance estimator (Conley 1999), which allows for correlation between areas that are geographically close but belong to different administrative units, using three different bandwidth choices and ignoring uncertainty in the first stage. The bandwidth choice has little impact on the standard error estimates, and the SHAC standard errors are very similar to estimates which allow for clustering on the district level.

In order to get an idea how ignoring uncertainty in the first stage affects standard error estimates, we report in column (5) standard errors clustered on the district level which do not take into account the uncertainty in the first stage. Estimates are not much smaller than the standard error estimates presented in column (1).

In column (6), we present standard errors clustered at the municipality level. Estimates are once again similar to those clustered at the district level.

D.3 Synthetic Cohort Results by Skill

In Figure O.1, we depict employment and wage gaps separately for unskilled and skilled workers between the highly exposed inner border region and their respective synthetic control (in bold).

To compare these outcomes with our baseline estimates, we must scale the differences in employment growth (about -0.086 for unskilled natives and -0.048 for skilled natives by 1993) and wage growth (about -0.016 for unskilled natives and -0.008 for skilled natives by 1993) by the share of Czech workers who entered the treatment region (5.8%). The results, about -1.49 (unskilled) and -0.82 (skilled) for employment, and -0.28 (unskilled) and -0.13 (skilled) for wages, are in the same ballpark as our baseline coefficients.

To test precision, we computed permutation tests in which we apply the synthetic control method to every potential control in our sample (as in Abadie, Diamond and Hainmueller, 2010). For unskilled workers, only one employment gap and five wage gaps are more negative than the treatment gap, indicating statistical significance at (at least) the 5% and 10% level. For skilled workers, 4 employment gaps are more negative than the treatment gap, indicating statistical significance at (at least) the 10% level. The treatment wage gap is however not exceptionally large in the treated inner border region compared to placebo districts. It should be noted that this permutation test is conservative, as the placebo districts are substantially smaller (and thus noisier) than the inner border region, which spans over multiple districts.

D.4 Employment Effects by Industry

Figure O.3 shows that industries which experienced a larger inflow of Czech workers experienced a larger decline in native local employment between 1990 and 1993. Specifically, we estimated employment equations separately for each of 28 industries and by skill, pooling over the years 1987 to 1993 to estimate linear municipality-specific time trends using the pre-shock period (as in column (2) in Table 5). The resulting slope coefficients reveal then the differential employment impact by industry. We then plot these coefficients against the industry's exposure to Czech workers, measured as the industry-specific share of Czech workers (relative to the mean share). The figures show that the employment decline in areas more exposed to Czech inflows tends to be concentrated in those industries in which the exposure to Czech immigrants is larger. Consistent with our analysis on the area level, this decline is larger among unskilled native workers. The slope of the corresponding regressions is equal to -1.78 (robust se .55) for unskilled and -.40 (se .26) for skilled workers. Reassuringly, employment declines are small in the public sector which hardly experienced any inflow in Czech workers and which may therefore be seen as a placebo check.

D.5 Impact on Firm Births and Native Job Creation

In the first two columns in Table O.3, we investigate whether the Czech inflow leads to an expansion of the number of firms in exposed areas. We report both unadjusted and trend-adjusted estimates and our sample includes the border region and matched control districts.

We indeed that that the inflow of Czech workers increases the number of firms. The effect is however relatively modest: a 10-percentage point increase in the employment share of Czechs increases the number of firms by about 3%. Interestingly, native job creation in newly created firms (measured as the number of native workers in new establishments in 1993 divided by native employment in 1990) declines in response to an immigration-induced labor supply shock, suggesting that many of the new hires in these firms are Czech workers (columns (3) and (4)). Interestingly, the expansion in the number of firms is entirely driven by the non-tradable sector, possibly because possibly because firm size and thus fixed entry costs are smaller in the non-tradable sector than in the tradable sector.

Overall, although the immigration-induced labor supply shock expanded the number of firms, the shock reduced native job creation in new firms. We wish to stress that these are the short-term effects of the labor supply shock. It may well be that in the longer run these new firms expand, creating positive employment possibilities also for native workers.

D.6 Wage and Employment Effects by Gender

In Table D.4, we report results separately for men and women, distinguishing also between skill groups. The sample includes the border region and matched control districts, and we display both unadjusted and trend-adjusted estimates.

Generally, the results show that the employment and wage effects are roughly similar for men and women. They also reveal that accounting for municipality-specific time trends has generally a stronger effect for women than for men, especially among skilled workers.

D.7 Alternative Specifications

In Table O.5, we compare employment and wage results from our baseline specifications with those obtained from specifications more commonly adopted in the literature. Specifically, we follow Borjas, Freeman and Katz (1996 AER P&P) and regress local native employment and wage growth of skilled and unskilled natives between 1990 and 1993 on the *skill-specific* change in the employment share of Czech workers (as opposed to the *overall* change, as in our baseline specification) in the area. The table has a similar structure as Table 2 in Borjas, Freeman and Katz (1996).

Panel A corresponds to our 2SLS estimates, where we predict the skill-specific inflow of Czech workers (i.e. the change in the number of Czech workers from a specific skill group in the municipality between 1993 and 1990 divided by skill-specific local employment in 1990) by a quadratic in distance to border, separately for skilled and unskilled workers. Without skill fixed effects in column (1), both the wage and employment coefficients are more negative than in our baseline specification. The reason is that the inflow of Czech workers was particularly high among the unskilled, who experienced lower employment and wage growth than skilled workers throughout Germany in the early 1990s. Controlling for skill fixed effects yields estimates that are closer to the baseline estimates reported in our manuscript (column (2)). Controlling for area fixed effects instead (column (3)) again yields more negative coefficient estimates, due to different secular trends for the two skill groups. If we include both area and skill fixed effects (column (4)), coefficient estimates are smaller than our baseline estimates, but still significantly negative.

To summarize: specifications (1) and (3) which do not include skill fixed effects yield biased estimates in our context, due to the differential secular wage and employment growth of the two skill groups. It is important to emphasize that specifications (2) and (4) (which control for skill fixed effects) identify conceptually a different parameter than our specification based on the overall inflow of Czech workers. In particular, specification (4) can identify only relative wage and employment effects between skill groups. Our paper suggests that in the short run both skill groups were negatively affected by Czech inflows (possibly because of the imperfect elasticity of capital), such that relative comparisons between skill groups underestimate the overall labor market impact of migration. Specification (2), which does not include area fixed effects, comes closest to out specification (and also yields similar results), but nevertheless identifies a different parameter. We prefer our specification, which uses *overall* rather than the skill-specific inflow of Czech workers for three main reasons. First, this specification consistent with our experiment, as only the total inflow of Czechs into the border region can be considered quasi-random. Second, the estimated parameters are clearly policy relevant in that it captures the total effect of the aggregate supply shock for specific groups of natives. And third, it avoids the problem of misclassification that arises when such observable characteristics are used to assign immigrants into skill groups in which they do not compete with natives.

For completeness, we report OLS estimates for specifications (1) to (4) in Panel B. Coefficient estimates are now much more positive in both the employment and wage regression. This suggests that the selection bias in migration destinations is large even if the outcome is differenced; that is, migrants systematically enter those areas which experience above-average wage and employment *growth*.

Table O.1: Alternative Functional Forms and Distance Measures for the First Stage

Panel A: Alternative Functional Forms					
Panel A	(1) Quadratic (Baseline)	(2) 3rd-order Polynomial	(3) Spline 5 knots	(4) Spline 10 knots	(5) Spline 20 knots
First Stage					
adj. R-sq	0.384	0.382	0.376	0.408	0.415
F	49.47	32.96	23.23	33.43	25.95
Second stage					
empl. Coef.	-0.947	-0.961	-0.961	-0.750	-0.652
	(0.380)	(0.380)	(0.380)	(0.390)	(0.365)
wage. Coef.	-0.124	-0.124	-0.127	-0.075	-0.102
	(0.103)	(0.103)	(0.102)	(0.104)	(0.091)
# municipalities	291	291	291	291	291

Panel B: Alternative Distance Definitions			
Panel B	(1) Airline distance (Baseline)	(2) Driving Distance	(3) Driving Time
First Stage			
adj. R-sq	0.384	0.387	0.347
F	49.47	48.45	40.02
Second stage			
empl. Coef.	-0.947	-0.838	-0.947
	(0.380)	(0.372)	(0.369)
wage. Coef.	-0.124	-0.118	-0.105
	(0.103)	(0.098)	(0.099)
# municipalities	291	291	291

Note: The table reports the adjusted R-squared and the F-statistic from alternative first stage regressions and the corresponding 2SLS estimates for the impact of the inflow of Czech workers into the municipality (measured as measured as the increase in the number of Czech workers between 1990 and 1992 as a share of employment in 1990) on native employment and log wage growth between 1990 and 1993 in the municipality. The sample is restricted to the border region. Panel A displays results for alternative functional form assumptions. Column (1) refers to our baseline estimates and uses a quadratic in airline distance to the nearest border crossing as instruments. Column (2) uses instead a 3rd-order polynomial, while Columns (3) to (5) use a spline of distance to border with 5, 10 or 20 knots. Panel B shows results for alternative measures of distance. Column (1) refers to our baseline estimates which use airline distance from the municipality's centroid to the nearest border crossing. Column (2) instead uses driving distance, while Column (3) uses driving times. Standard errors clustered at the district level in parentheses.

Data Source: German Social Security Records, border region, 1993.

Table O.2: Alternative Computations of Standard Errors

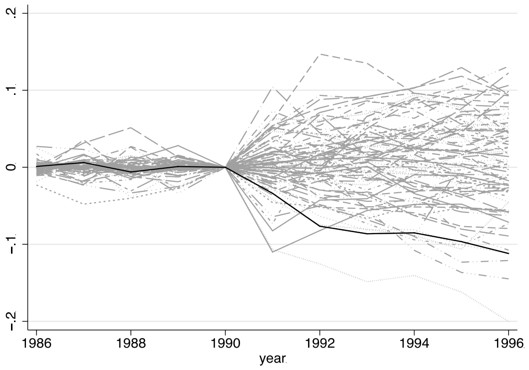
Panel A: Wage Effects			SHAC		Clustering	
	1	2	3	4	5	6
	district	bw=50km	bw=100km	bw=200km	district, ignore first stage	municipality
all	0.002	0.002	0.002	0.002	0.002	0.002
	(0.066)	(0.0502)	(0.0530)	(0.0643)	(0.062)	(0.051)
unskilled	-0.057	-0.057	-0.057	-0.057	-0.057	-0.057
	(0.106)	(0.0820)	(0.0800)	(0.0851)	(0.099)	(0.086)
skilled	-0.052	-0.052	-0.052	-0.052	-0.052	-0.052
	(0.057)	(0.0506)	(0.0500)	(0.0565)	(0.056)	(0.049)
Panel B: Employment Effects			SHAC		Clustering	
	1	2	3	4	5	6
	district	bw=50km	bw=100km	bw=200km	district, ignore first stage	municipality
all	-0.930	-0.930	-0.930	-0.930	-0.930	-0.930
	(0.243)	(0.236)	(0.243)	(0.230)	(0.210)	(0.263)
unskilled	-1.203	-1.203	-1.203	-1.203	-1.203	-1.203
	(0.328)	(0.275)	(0.271)	(0.263)	(0.261)	(0.293)
skilled	-0.522	-0.522	-0.522	-0.522	-0.522	-0.522
	(0.214)	(0.226)	(0.230)	(0.213)	(0.206)	(0.243)
# municipalities	1,550	1,550	1,550	1,550	1,550	1,550

Note: The table compares standard errors using alternative estimatin procedures. The sample includes both the border region and matched inland control districts and the specification refers to the long difference regression (as in column (v) in Table 5). Columns (1) to (3) report standard error estimates based upon the spatial HAC technique of Conley (1999), using a uniform kernel and bandwidths of 50, 100 or 200 kilometres, respectively. Columns (4) to (6) report standard errors from a wild bootstrap procedure. While column (4) ignores uncertainty in the first stage, Columns (5) and (6) do not. Columns (4) and (6) cluster on the district, Column (5) on the municipality level.
Data Source: German Social Security Records, border region and matched control districts, 1990 and 1993.

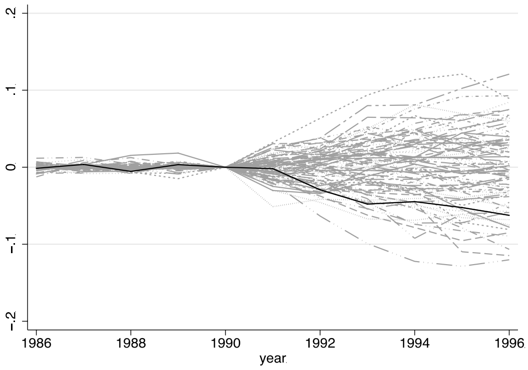
Figure O.1: Synthetic Control Method for Unskilled and Skilled Native Workers, with Permutation Tests

Panel A: Employment

A.1 Unskilled

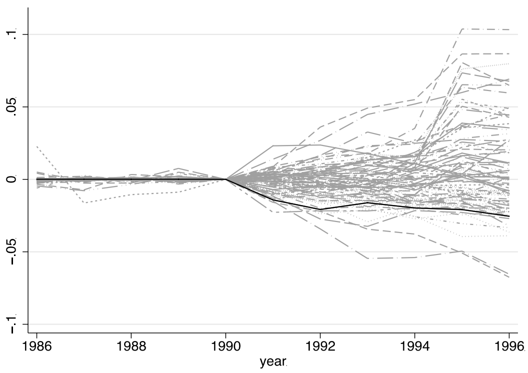


A.2 Skilled

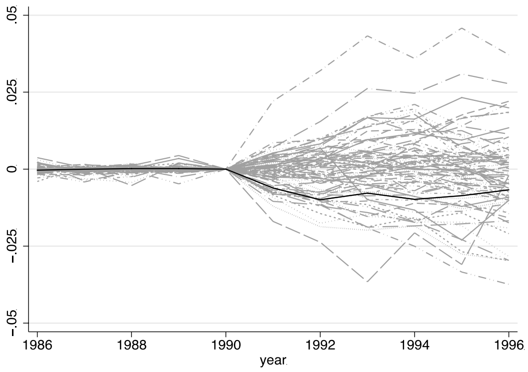


Panel B: Wages

B.1 Unskilled



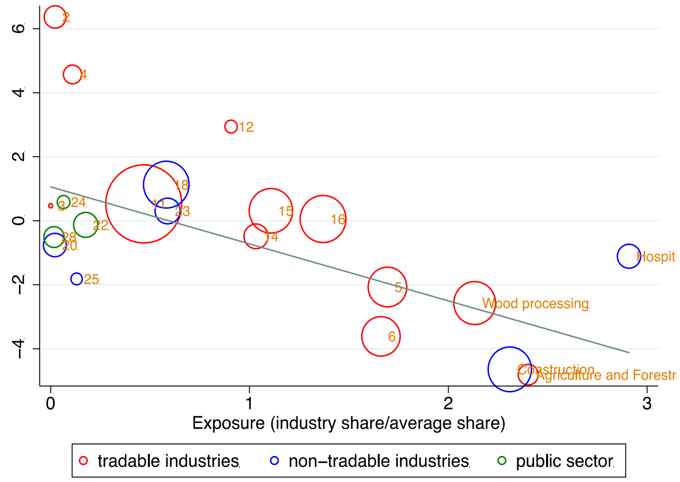
B.2 Skilled



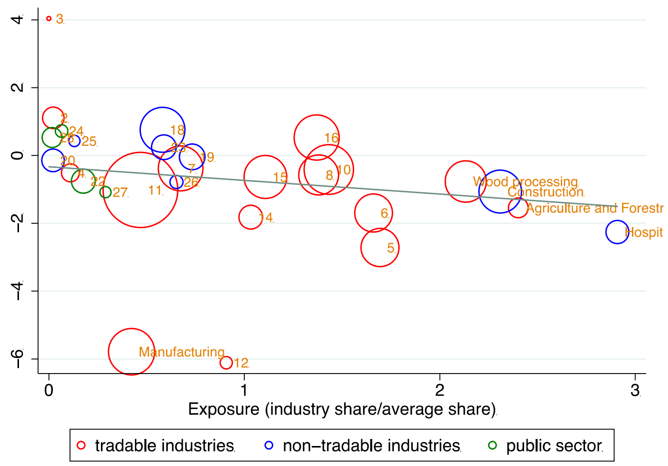
Note: The figures display, following Abadie et al. (2010), the employment and wage gaps between the highly affected inner border region and its synthetic control (bold line) as well as the respective gaps from 85 placebo districts, separately for unskilled and skilled workers. To make sure that the wage effects are not underestimated because of worker selection, wage growth is computed for all workers who remain employed in the district between two consecutive years, as in our baseline specification.
Data Source: German Social Security Records, inner border region and inland districts, 1986-1996.

Figure O.2: Employment Effects by Industry

Panel A: Unskilled



Panel B: Skilled



Industry codes:
[1] Agriculture and forestry
[2] Energy
[3] Mining
[4] Chemical industry
[5] Plastics
[6] Pit and quarry
[7] Ceramic and glass
[8] Metal production and processing
[9] Manufacturing
[10] Vehicle manufacturing
[11] IT, electronics, optics
[12] Musical instruments, jewelry, toys
[13] Wood and wood processing
[14] Printing and paper processing
[15] Leather and textile
[16] Food and tobacco
[17] Construction
[18] Trading
[19] Transportation and communications
[20] Credit and insurance
[21] Hospitality
[22] Healthcare and welfare
[23] Business-related services
[24] Educational services
[25] Recreational services
[26] Household services
[27] Social services
[28] Public administration

Note: The Figure plots the coefficients from a regression of employment growth of unskilled (Panel A) or skilled (Panel B) native workers in the industry on the predicted (by the square in distance to border) local overall inflow of Czech workers against the ratio between the industry-specific and average share of Czech workers. Employment growth regressions are pooled over the years 1987 to 1993 to estimate municipality-specific linear time trends using the pre-shock period. The size of each circle is proportional to industry-specific employment in 1990. See side legend for industry codes.
Data Source: German Social Security Records, border region and matched control districts, 1987 to 1993.

Table O.3: Impact of Czech Inflows on Number of Firms and Native Job Creation

	# of establishments		native job creation by establishment entry		# of establishments in tradable sector		# of establishments in non-tradable sector	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
all	0.356	0.305	-0.248	-0.197	-0.029	0.097	0.413	0.440
	(0.113)	(0.171)	(0.101)	(0.154)	(0.197)	(0.262)	(0.158)	(0.201)
unskilled	-	-	-0.467	-0.354	-	-	-	-
			(0.139)	(0.198)				
skilled	-	-	-0.192	-0.194	-	-	-	-
			(0.099)	(0.157)				
trend controls	no	yes	no	yes	no	yes	no	yes
# municipalities	1,550	1,550	1,550	1,550	1,550	1,550	1,550	1,550

Note: Columns (1) and (2) show 2SLS estimates for the impact of the inflow of Czech workers into the municipality (measured as the increase in the number of Czech workers between 1990 and 1992 as a share of employment in 1990 and instrumented by a quadratic in airline distance to border) on the percentage change in the number of establishments in the municipality between 1990 and 1993. Column (1) refers to the period 1990-1993, while Column (2) pools over the period 1987-1993 to estimate municipality-specific linear trends using the pre-shock period. Columns (3) and (4) show the corresponding estimates for native job creation, measured as the number of all (unskilled, skilled) native workers in new establishments in 1993 divided by native (unskilled, skilled) employment in 1990. Columns (5) to (8) display 2SLS estimates for the impact of the inflow of Czech workers into the municipality on the percentage change in the number of establishments in the municipality between 1990 and 1993 separately for the tradable and non-tradable sector. Bootstrapped standard errors clustered on the district level are reported in
Data Source: German Social Security Records, border region and matched control districts, 1987 to 1993.

Table O.4: Impact of Czech Inflows on Male and Female Employment and Wage Growth

	unadjusted		trend-adjused	
	(1)	(2)	(1)	(2)
	men	women	men	women
Panel A: Wage Effects				
all	-0.114	-0.198	-0.151	-0.320
	(0.043)	(0.054)	(0.060)	(0.064)
unskilled	-0.233	-0.233	-0.187	-0.396
	(0.067)	(0.091)	(0.066)	(0.111)
skilled	-0.085	-0.072	-0.147	-0.265
	(0.043)	(0.056)	(0.062)	(0.064)
Panel B: Employment Effects				
all	-0.974	-0.851	-0.955	-0.902
	(0.240)	(0.246)	(0.378)	(0.330)
unskilled	-1.871	-1.038	-1.540	-1.367
	(0.308)	(0.362)	(0.563)	(0.477)
skilled	-0.680	-0.086	-0.916	-0.835
	(0.244)	(0.233)	(0.369)	(0.316)
# municipalities	1,550	1,550	1,550	1,550

Note: The table presents 2SLS estimates for the impact of the inflow of Czech workers into the municipality (measured as the change in the number of Czech workers between 1992 and 1992 divided by employment in 1990 and instrumented by a quadratic in airline distance to border) on local native wage and employment growth between 1993 and 1990, separately for men and women. We report both unadjusted estimates (columns (1) and (2)) and trend-adjusted estimates which use the 1987-1989 pre-shock period to estimate municipality-specific linear time trends. Bootstrapped standard errors clustered on district level in parentheses.
Data Source: German Social Security Records, border region and matched control districts, 1987 to 1993.

Table O.5: Impact of the Skill-Specific Inflow of Czech Commuters on Skill-Specific Native Wage Growth, 1993-1990

Panel A: 2SLS	(1)	(2)	(3)	(4)
Wages	-0.228	-0.163	-0.249	-0.106
	(0.047)	(0.092)	(0.043)	(0.060)
Employment	-2.134	-0.612	-2.985	-0.127
	(0.263)	(0.317)	(0.338)	(0.397)
Skill fixed effects	no	yes	no	yes
Municipality fixed effects	no	no	yes	yes

Panel B: OLS	(1)	(2)	(3)	(4)
Wages	-0.086	-0.032	-0.097	-0.004
	(0.024)	(0.033)	(0.037)	(0.038)
Employment	-0.700	0.043	-1.240	-0.063
	(0.095)	(0.128)	(0.251)	(0.105)
Skill fixed effects	no	yes	no	yes
Municipality fixed effects	no	no	yes	yes
N: 582 (291 municipalities X 2 skill groups)				

Note: The table reports 2SLS (Panel A) and OLS (Panel B) estimates of the skill-specific inflow of Czech workers into the municipality (measured as the change in the number of Czech workers of a particular skill-group divided by skill-specific employment in 1990) on native employment and log wage growth in the municipality. In Panel A, the skill-specific inflow of Czech workers is instrumented with a quadratic in airline distance to the border. In Column (1), neither skill nor municipality fixed effects are included. In Columns (2) and (3), either skill or municipality fixed effects are included. In Column (4), both skill and municipality fixed effects are controlled for. Standard errors clustered on the district level are reported in parentheses.
Data Source: German Social Security Records, border region, 1990 to 1993.