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Trade and towns: Heterogeneous adjustment to a border shock☆

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ABSTRACT

We study the effects of changes in trade openness on wages and employment across towns of different sizes, using Austrian regional data and the fall of the Iron Curtain as a quasi-experimental setting. We find improved access to foreign markets to boost both employment and nominal wages in border regions, but large towns tend to have larger wage responses and smaller employment responses than small towns. These adjustment patterns are consistent with a multi-region model featuring labor supply elasticities that vary with town size. The implied differential border-town welfare gains are related non-monotonically to town size, peaking at a population level of about 150,000.

1. Introduction

We estimate the effects of trade liberalization on employment and wage growth of different-sized towns within a country. To this end, we track the evolution of employment and wages in fine-grained regional data for Austria, arguing that the fall of the Iron Curtain in 1990 represents a large and fully exogenous trade shock to the Austrian economy. We define eastern border regions as the 'treatment group' and the rest of Austria as the 'control group.' Hence, the three differences are (i) before vs after 1990, (ii) border vs interior towns, and (iii) large vs small border towns. Access to foreign markets is found to boost both factor quantities and prices, as wages and employment on average grew more strongly post-1990 in treatment regions than in control regions. However, we find significant heterogeneity in these responses across the size distribution of towns. Larger towns are characterized by larger nominal wage responses and smaller employment responses than smaller towns. Hence, local labor supply is found to be less elastic in large towns than in small towns.

These findings are in line with the predictions of a multi-region model of intra-national adjustment to trade. In our model, towns are heterogeneous in their exposure to trade and in their relative endowments of fixed and mobile factors. Intra-national adjustment takes place via labor migration: workers move in search of the highest real wage, with immobile housing and heterogeneous locational preferences acting as dispersion forces. The model sets up our structural triple-difference empirical strategy and predicts that trade liberalization will trigger a stronger wage increase but weaker employment increase in larger towns.

A potential alternative mechanism consistent with our empirical findings is skill heterogeneity, e.g. with border towns hosting more skill-intensive sectors than interior towns, or with bigger towns being populated by more skilled workers. We test for effects related to sector and skill compositions by matching border to interior towns based on industrial structure and by focusing on blue collar workers only. Our qualitative results are robust to these as well as to a number of additional variations on the baseline empirical model.

Heterogeneous spatial effects of openness shocks have been researched carefully before, but our paper innovates on four main counts. First, we estimate wage as well as employment effects of such shocks. Employment responses are relatively well understood following the seminal paper on the effects of German division and reunification by Redding and Sturm (2008) and subsequent research, but the Austrian

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data are unique in containing also wages at a fine level of spatial disaggregation. We establish that wages respond to changes in trade exposure qualitatively differently from employment. Second, while German episodes of partition and re-unification implied joint trade and migration shocks, the experience of Austria pre- and post-1990 is much closer to a pure trade shock, as trade was liberalized swiftly after the collapse of the socialist regimes but cross-border migration continued to be tightly controlled well into the early 2000s. Third, the existing literature focuses on the average effect of increased trade exposure. We shift our attention to the heterogeneous response across different-sized towns following the trade shock. Finally, our structural approach enables us to estimate differential welfare effects of town-level exposure to international trade. The model suggests that places that add the greatest absolute number of jobs are those that also gain the most in terms of the average wellbeing of their incumbent residents. As large towns experience relatively large wage swings and relatively small adjustments in terms of employment, the welfare effects on inframarginal workers (those who do not move) are related non-monotonically to town size.

Our results are also relevant to economic policy. Policy makers commonly expect international trade to benefit urbanized regions more than rural regions.[1] For this reason, trade reforms are often accompanied by transfer schemes designed to compensate rural regions.[2] We show below that international trade appears to favor relative employment growth in smaller towns disproportionally. In terms of labor quantities, therefore, trade would seem to promote spatial convergence. We argue that such an analysis falls short as it fails to consider factor price effects, and we find that large towns experience proportionally larger wage gains, offsetting the equalizing patterns of employment growth. Taken literally, our model and estimations imply the biggest trade gains for incumbent inhabitants of medium-sized towns with a population of around 150,000. This might offer an explanation for the apparent contradiction between prior empirical findings and the predominant view held by policy makers.

Our paper is organized as follows. In Section 2, we offer a brief review of the literature. Section 3 presents our triple-difference estimation strategy, our empirical setting and our estimation strategy. The empirical results are reported in Section 4. Section 5 presents a multi-town model of spatial adjustment allowing us to put structure on the empirical analysis. We use the model and the empirical estimates for some welfare calculations in Section 6. Section 7 concludes.

2. Literature background

2.1. Theory

The question we address can be formulated as follows: does a given change in external market access affect employment, wages and welfare differently in small and large towns? We can distinguish essentially two theoretical approaches to this question.

One approach focuses on differences between small and large towns in terms of economic self sufficiency. Their very size allows large towns to produce a larger range of differentiated goods and hence to be economically more self-contained, whereas small towns are comparatively more open to trade with the rest of the economy – analogous to the impact of country size on trade openness as represented in the gravity model of international trade. Thus, a given reduction in international trade costs will have a bigger impact on small towns, as they rely more on trade than large towns. This is the mechanism emphasized by Redding and Sturm (2008) and developed by Helpman (1998). Importantly, as long as local labor supply is neither perfectly elastic nor perfectly inelastic, this mechanism implies that, after trade liberalization, small towns will experience stronger increases in both employment and nominal wages than large towns.[3] However, this positive correlation between employment and wage changes is rejected by our estimations, which strongly point towards employment effects falling with town size while wage effects rise with town size. An alternative mechanism therefore seems to be at play, at least in our data.

The obvious alternative approach is to focus on differences in factor supply elasticities between small and large towns. Combes et al. (2005) capture the essence of this effect through graphical analysis. By considering the possibility of imperfectly elastic local labor supply they highlight the importance of local supply conditions in determining the wage and employment effects of a given trade shock. In contrast to models relying on a gravity-type mechanism, this approach opens up the possibility that employment and nominal wages respond differentially across different types of towns. Monte, Redding and Rossi-Hansberg (2016) develop a general-equilibrium model of differing local labor supply elasticities based on differing potential commuting flows. Using U.S. county-level data, they find no evidence of a systematic reduced-form link between own-county employment and the local labor supply elasticity. It has, however, been shown that housing supply is more elastic in areas that are less dense and therefore have more available land (Hilber and Mayer, 2009; Saiz, 2010; Hsieh and Moretti, 2017). If housing supply is less elastic in larger, denser towns, then this would plausibly lead to larger towns having less elastic labor supplies as well. Our model formalizes such a mechanism via non-homothetic preferences for locational amenities.

It might be useful to make explicit what we do not consider. First, we abstract from exogenously determined comparative advantage across towns and countries (though we shall take this possibility into account in the empirics). In Henderson (1982), for example, trade liberalization is found to increase the number of towns that are specialized in the comparative advantage goods. Since towns specialized in capital-intensive sectors are bigger in equilibrium, trade liberalization will favor the growth of larger towns in capital abundant countries and of smaller towns in capital scarce countries. In this model, welfare is equalized across towns. In a similar vein, Autor et al. (2013), Monte (2016) and Dix-Carneiro and Kovak (2017) explore the impact of growing import penetration respectively in the United States and Brazil, taking initial industry specialisation as the regional trade exposure measure and abstracting from town size. Second, we abstract from the differential intensity of the liberalization shock across locations (except in the form of our distinction between treatment and control regions), and we focus on differential responses to a pure trade shock of given intensity. This contrasts with a recent literature on responses to external opening of goods and factor markets, considering intra-national spatial frictions (e.g. Atkin and Donaldson, 2015; Cosar and Fajgelbaum, 2016; Fajgelbaum and Redding, 2014; Ramondo et al., 2016), and on the differential impact across different-sized towns of improved market access through better transport infrastructure (e.g. Baum-Snow et al., 2017). Unlike Baum-Snow et al. (2017), we cannot distinguish between domestic and international market access, and instead focus on the latter only. Finally, our empirical setting does not allow us to distinguish between trade liberalization in intermediate and final goods, which can have important distributional effects, though not necessarily across regions (De Loecker et al., 2016).

2.2. Empirics

Existing empirical work on the city-level effects of trade opening

[1] According to the World Bank (2008; p. 12), for instance, "openness to trade [...] makes subnational disparities in income larger and persist for longer. [...] Economically dense places do better."

[2] The European Union's regional policy is the best known example. Conditional cash transfer (CCT) schemes in developing countries are often motivated by trade reforms and typically targeted at rural households (Fiszbein and Schady, 2009).

[3] The same qualitative result is found, among others, in Krugman and Livas Elizondo (1996) and Behrens et al. (2007).

exclusively focuses on employment or population. This evidence strongly points toward spatially equalizing effects of trade.

Cross-country panel regressions suggest that trade reduces urban primacy (see e.g. Ades and Glaeser, 1995; Henderson, 2003), but measurement and identification are challenging at that level. Redding and Sturm (2008) identify causal effects by focusing on the quasi-experimental setting offered by post-War German partition and subsequent national reunification. Of particular interest to us is the distinction they draw between initially larger and smaller West German border cities (the treatment sample), and their observation that German partition had a more severe impact on population growth of smaller cities than on comparable larger cities. This result corroborates the central finding from the cross-country literature: access to foreign markets disproportionately promotes the growth of smaller cities.

Due to data limitations, these studies could not track the effect of trade liberalization on city-level wages. Ahlfeldt et al. (2015) explore spatial effects of the Iron Curtain using data on employment, wages, land prices and travel times. Their analysis, however, focuses on adjustment at the intra-city level (within Berlin). Hanson (1997) and Kovak (2013) estimate effects of trade liberalization on regional wages, but their papers do not differentiate regions by size, density or urbanization. Simultaneous wage and employment effects are estimated in a precursor paper Brülhart et al. (2012), but that paper does not investigate the heterogeneous effects that interest us here, nor does it offer a theory that leads formally to the empirical model.

To the best of our knowledge, this is the first study to analyze jointly what happens to wages and employment across different-sized towns as external trade is liberalized.

3. Empirical setting and estimation strategy

3.1. Austria and the end of the Iron Curtain: a case of exogenous trade liberalization

Austria offers a propitious setting, akin to a natural experiment, within which to explore regional responses to changes in trade openness. Austria has long been a very open economy, with exports and imports corresponding to 58% of GDP in 1975 and to 93% in 2002. It was the OECD's fifth most trade-oriented country in 1990. Despite its geographic centrality, however, post-War Austria had lain at Europe's economic periphery for more than four decades. In 1976, at the beginning of our sample period, the country still belonged to the eastern edge of democratic, market-oriented Europe. By 2002, which marks the end of our sample period, it had become the geographical heart of a continent-wide market economy. Crucially for our study, the fall of the Iron Curtain in the second half of 1989 triggered a change in trade openness that was large and unanticipated. Similarly important, during the period covered by our study, this transformation took the form of an almost pure trade shock: a large change in cross-border openness of goods markets associated with continuing segmentation of cross-border labor markets. As a consequence, by 2002 Austrian trade shares with the country's eastern neighbors had more than doubled, while those with its established western neighbors had shrunk by up to 20% . This increase was significant in absolute terms as well: the value of Austria's trade with its eastern neighbors increased from 2.9 to 7.4% of GDP over our sample period - much faster than the increase in trade with its western neighbors from 31.0 to 33.2% of GDP.[4]

We define 1990 as the moment that marked the general recognition of a lasting economic transformation of the Central and Eastern European countries (CEECs) and of their new potential as trade partners. Actual trade barriers, however, did not fall immediately. Hence, the decade following 1990 was a period of gradual but profound and lasting mutual opening of trade, to an extent that had been largely unanticipated up to the very late 1980s. Austria's east-west elongated geography accentuates the fact that access to the eastern markets becomes relatively less important than access to western markets as one crosses Austria from east to west. This offers us the required identifying variation for the estimation of trade effects. We compare post- versus pre-1990 trends in eastern Austrian border regions (the 'treatment group') with post- versus pre-1990 trends in the rest of Austria (the 'control group') as well as with the western border regions (the 'placebo group'). To the extent that no other major exogenous change affected the treatment group over the treatment period, the resulting difference-in-difference estimates can be interpreted as the causal effects of increased trade openness.

3.2. Data

Our key variables are municipality-level employment and wages computed from the Austrian Social Security Database (ASSD). The ASSD reports individual labor-market histories, including wages, for the universe of Austrian workers.[5] These records can be matched to establishments, which allows us to allocate workers to locations. We observe wages and employment at three-month intervals, taken at the mid point of each quarter, yielding 108 measurements from the first quarter of 1976 to the fourth quarter of 2002.

The ASSD assigns every establishment to one of 2305 municipalities. We treat municipalities as individual locations unless they are included in one of the 33 functional urban areas defined by Statistics Austria, in which case they are aggregated as one location. Our "towns" therefore are either a (mostly small) municipality or a group of municipalities that is defined by the statistical office as an integrated urban area. Our data set contains 2047 towns. In order to minimize distortions from top coding, we construct wages as medians across individuals by town. Wages are recorded on a per-day basis, which means that they are comparable irrespective of whether employment contracts are part-time or full-time. Table 1 provides descriptive statistics.

Our identification strategy will hinge on the relative distances of these towns to eastern markets. We retain two distance measures, both based on road distances in order to account for the mountainous topography of much of Austria:

- the road distance to the nearest border crossing with a CEEC country (see Fig. 1),
- the road distance to the nearest CEEC town with a population of at least 50,000 or 20,000 in 1990 (see Fig. 2).[6]

Fig. 3 illustrates the key relation we exploit for our empirical analysis, by showing the estimated post-1990 growth differential of town-level wage bills against the towns' distance from the eastern border based on natural spline regressions.[7] The plot shows that there is a statistically significantly positive wage-bill effect for municipalities that

[4] The transition toward free cross-border worker mobility began with EU enlargement in 2004. Throughout our post-liberalization sample period 1990–2002, however, integration between the CEECs and the EU was characterized by very limited labor flows but strong trade integration and increasing capital market integration (OECD, 2001). For additional institutional and statistical details on the trade shock implied for Austria by the fall of the Iron Curtain, as well as for evidence on the intra-national spatial gradient of economic links to neighboring countries, see Brülhart et al. (2012).

[5] For a full description, see Zweimüller et al. (2009). Due to missing data for public-sector workers and the self-employed, we work exclusively with data pertaining to private-sector employees.

[6] Road distances were obtained from Digital Data Services GmbH, Karlsruhe, Germany. Only border crossings allowing for the handling of trucks carrying 3.5 tons or more are considered. These data pertain to measurements taken in the early 1990s. While some cross-border roads have been upgraded after 1990, we are not aware of any significant new border crossings that have been constructed between 1990 and 2002, except for a highway link with Slovenia that was opened in 1991.

[7] The smoothed line is obtained by estimating a cubic polynomial of y, the post-versus-pre-1990 growth differential of town-level wage bills, against x, distance from the eastern border. This estimation is performed between every pair of nodes under the constraint of continuity at each of the seven nodes.

Table 1
Summary statistics
Border defined as lying within 25 road kilometers from nearest CEEC border crossing (Czech Republic, Slovakia, Hungary or Slovenia).

Variables	1976–1989				1990–2002			
	Mean	Std. Dev.	Min	Max	Mean	Std. Dev.	Min	Max
Border municipalities (Border = 1)	16,484 observations: 52 quarters and 317 towns				17,752 observations: 56 quarters and 317 towns			
Median daily wage (Austrian Schillings)	310.87	92.38	53.10	765.11	608.34	139.88	190.08	1,374.48
Annual growth rate of median wage %, ΔW	6.2864	9.73	−103.07	128.01	4.1742	8.88	−100.57	81.08
Employment	502.21	3,492.11	1	67,239	563.15	3,798.11	1	71,739
Annual growth rate of employment %, ΔE	1.1168	14.07	−163.19	165.43	1.8767	21.38	−178.43	185.90
Minimum road distance to Eastern border (km)	14.90	6.97	0.00	24.91	14.90	6.97	0.00	24.91
Minimum road travel time to Eastern border (min)	21.29	9.71	0.00	48.17	21.29	9.71	0.00	48.17
Interior municipalities (Border = 0)	89,960 observations: 52 quarters and 1730 towns				96,880 observations: 56 quarters and 1730 towns			
Median daily wage (Austrian Schillings)	337.22	100.72	35.52	950.62	626.61	144.00	111.65	1,660.97
Annual growth rate of median wage %, ΔW	5.9667	9.46	−135.23	150.66	3.6581	8.85	−121.18	125.14
Employment	1,071.09	15,152.95	1	617,433	1,132.63	15,519.61	1	642,011
Annual growth rate of employment %, ΔE	1.2693	14.75	−179.86	181.54	1.2289	19.81	−185.38	184.62
Minimum road distance to Eastern border (km)	128.81	121.48	25.01	523.00	128.81	121.48	25.01	523.00
Minimum road travel time to Eastern border (minutes)	109.76	76.73	17.67	373.33	109.76	76.73	17.67	373.33

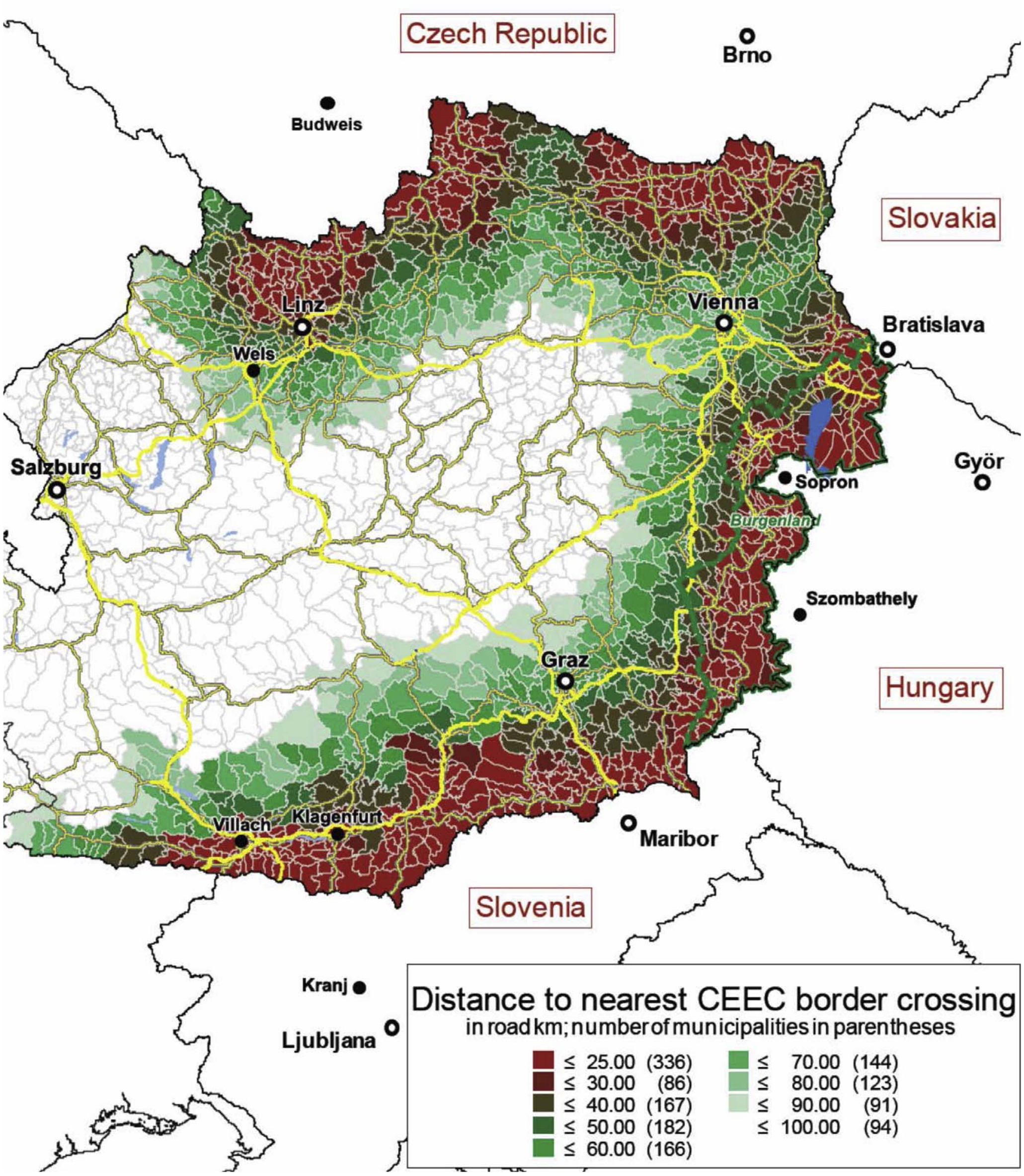

Fig. 1. Treatment groups: baseline definition.

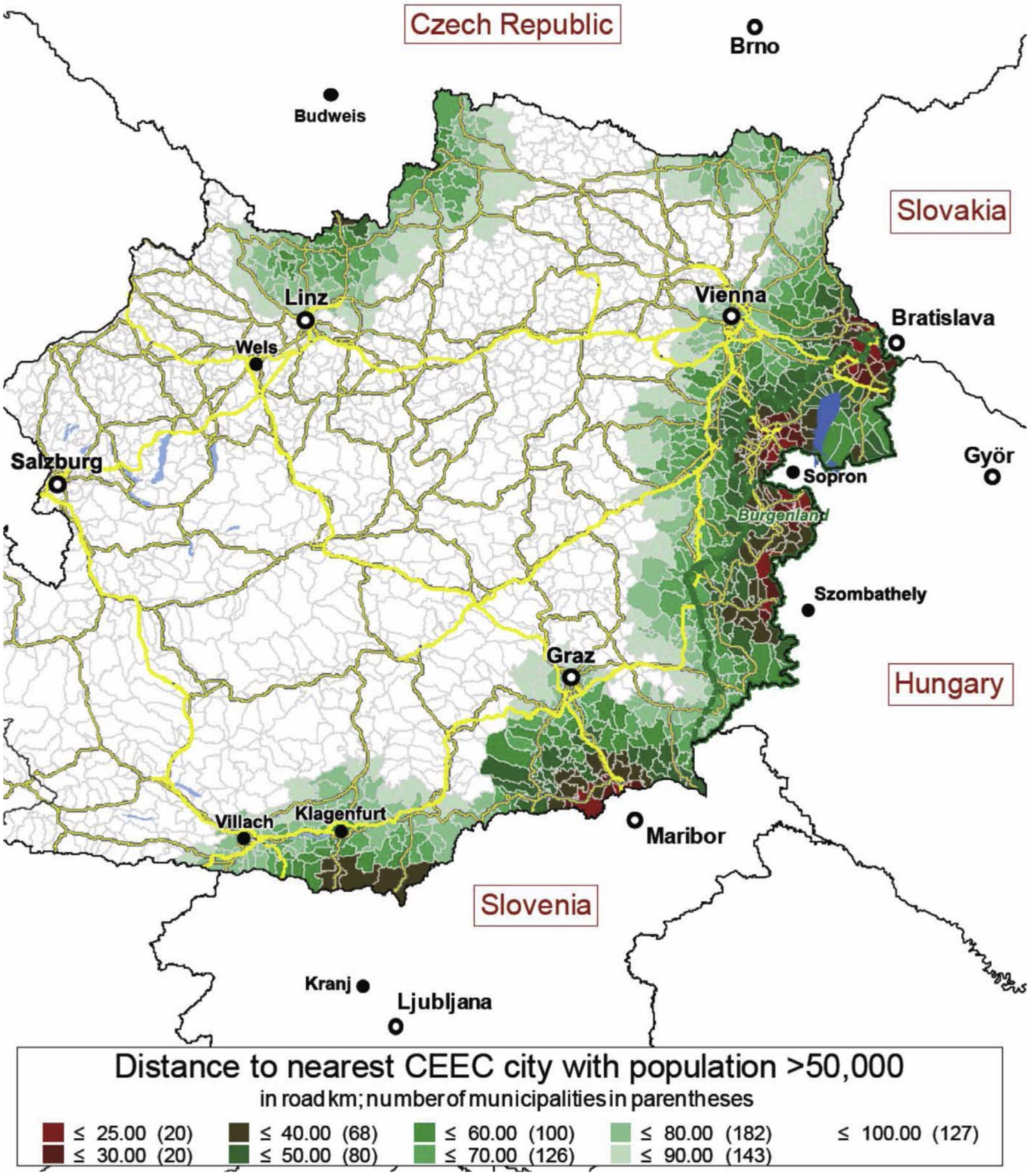

Fig. 2. Treatment groups: alternative definition.

are located close to Austria's eastern border, whereas there is none for municipalities beyond about 70 km from the border, with Vienna, at 65 km, still significantly affected. The differential effect of post-1990 market opening was thus confined to a relatively narrow band of towns located close to the border.

3.3. Estimation strategy

We exploit our quasi-experimental set-up for the following triple-difference estimation:

$$\Delta(W_{jt} \times E_{jt}) = \alpha_1(Fall_t \times Border_j \times Size_j) + \alpha_2(Fall_t \times Border_j) + \alpha_3(Fall_t \times Size_j) + d_j + d_t + \varepsilon_{jt}^{WE}, \quad (1)$$

$$\Delta W_{jt} = \beta_1(Fall_t \times Border_j \times Size_j) + \beta_2(Fall_t \times Border_j) + \beta_3(Fall_t \times Size_j) + d_j + d_t + \varepsilon_{jt}^{W}, \quad (2)$$

$$\Delta E_{jt} = \gamma_1(Fall_t \times Border_j \times Size_j) + \gamma_2(Fall_t \times Border_j) + \gamma_3(Fall_t \times Size_j) + d_j + d_t + \varepsilon_{jt}^{E}, \quad (3)$$

where W_{jt} is the nominal wage in town j and period t; E_{jt} is employment; $Size_j$ denotes mean-differenced town-level employment averaged over the pre-treatment period 1976–1989 in units of 10,000; $Border_j$ is a dummy for border (i.e. treatment) regions; $Fall_t$ is a dummy for quarters from 1990 onwards (the treatment period); d_j and d_t are town and time fixed effects, respectively; and ε_{jt}^{WE}, ε_{jt}^{W} and ε_{jt}^{E} are stochastic error terms.[8] Δ denotes year-on-year percentage changes.[9] Hence, unobserved time-invariant heterogeneity in town-specific wage and employment levels are differenced out. Moreover, the town dummies control for any unexplained differences in linear trends, and the time dummies control for nation-wide temporary shocks to wage and employment growth, including the common impact of the fall of the Iron Curtain on median

[8] For a theoretical underpinning to this triple-difference empirical model, see Section 5.

[9] Specifically, $\Delta X_{jt} \equiv \frac{X_{jt} - X_{jt-4}}{[X_{jt} + X_{jt-4}]\frac{1}{2}}$, for $X \in \{E, W, W \times E\}$.

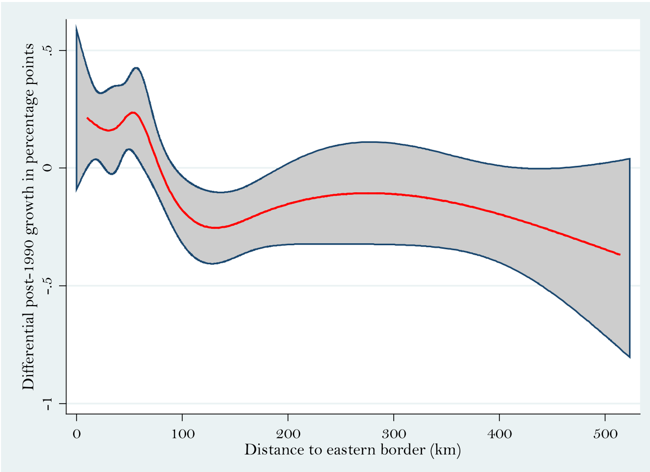

Fig. 3. Distance to border and post-1990 growth of town-level wage bills.
Notes: The graph reports estimates from a spline regression of post-1990 differential growth of town-level wage bills on towns distance from the eastern border. Sample: 1976-2002; 2,047 towns.

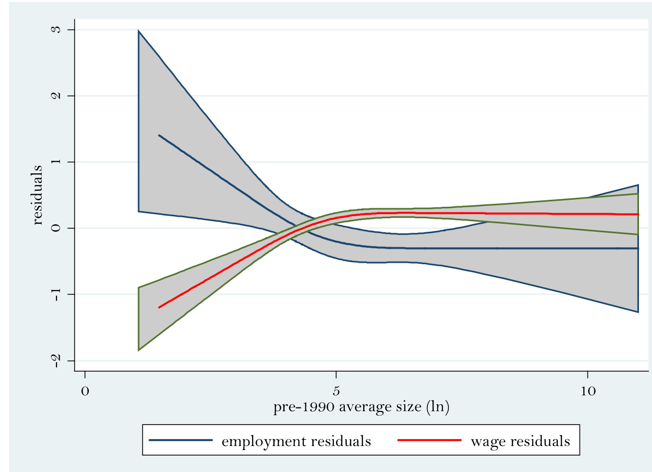

Fig. 4. Wage/employment responses and town size.
Notes: The graph reports estimates from a spline regression of residuals of Eqs. (2) and (3) (not including the triple interaction) on the log of pre-1990 town-level employment. Sample: 1976–2002; 2047 towns.

wage and employment growth across all of Austria.

In an alternative specification, we seek to control for the possibility that border regions differ systematically from interior regions not only in terms of proximity to the border but also in terms of size and industrial composition. We therefore reduce the set of control (i.e. interior) municipalities to those that provide the nearest match to at least one of the treatment (i.e. border) municipalities in terms of the sum of squared differences in sectoral employment levels, measured in 1989. We compute parameter estimates as average treatment effects in big and small towns where we match municipality-specific differential pre-versus-post-1990 growth rates between pairs of border and interior municipalities with the most similar sectoral employment structures.

Standard errors are clustered by municipality in all of our estimations.

4. Estimation results

4.1. Baseline estimates

We begin with a non-parametric illustration of our central finding. In Fig. 4, we plot residuals of Eqs. (2) and (3), not including the triple interaction terms, against the log of pre-1990 town-level employment. It becomes clearly apparent that small towns have stronger employment effects and weaker wage effects than large towns, and that this configuration reverses as one moves up the distribution of town sizes. This finding ermerges consistently also across our parametric estimates. Our baseline results are shown in Table 2. We report estimates of Eqs. (1) to (3) for four different definitions of $Border_j$, our indicator variable for the treatment sample. The coefficients on the interaction term ($Fall_t \times Border_j$) are positive throughout and mostly statistically significant. This shows that, compared to interior towns, towns close to Austria's eastern border have experienced stronger growth in both employment and wages after the fall of the Iron Curtain. This effect, however, was unevenly shared across border towns. Our estimated coefficients on the triple interaction ($Fall_t \times Border_j \times Size_j$) are consistently positive and statistically significant for wage changes and negative and statistically significant for employment changes. Thus, larger towns seem to have responded to external trade opening mainly through wage rises, whereas small towns responded mainly through employment growth. The effect on the total municipal wage bill, however, seems not to differ systematically between small and large towns, the point estimate on the triple interaction being statistically indistinguishable from zero in our baseline specification. In Section 5, we shall present a model that can accommodate all these qualitative adjustment patterns.

Specifications (2) and (3) allow us to estimate effects averaged over the full 1990–2002 treatment period. It is straightforward to document the timing of adjustment by estimating effects separately for each year through the inclusion of interacted year dummies, separately for small and large towns. We illustrate these effects in Figs. 5 and 6. The graphs show very similar patterns prior to 1990, but clearly above-average subsequent growth in large-town wages and in small-town employment. These graphs show furthermore that our chosen treatment period is long enough: by 2002 employment and wages no longer grew disproportionally in border towns. Finally, we observe that wage effects were strongest in the 1995–1998 period, whereas the employment growth of small towns peaked in the 1997–2001 period. Wage effects therefore preceded employment effects by some two years.

4.2. Robustness

4.2.1. Confounding factors

Our aim is to identify the causal effect of trade openness on municipal labor markets. We therefore subject the baseline estimates to a range of sensitivity tests that take accound of potential confounding factors and measurement issues. First, we add a dummy variable for the state of Burgenland post-1995, as this economically lagging region became eligible for generous EU subsidies after Austria joined the EU in 1995 and could thereby drive our estimated treatment effects. Second, we estimate the baseline model without including Vienna, to control for a potentially distorting effect of urban primacy in the control group.[10] Third, we truncate our sample at the other end of the size distribution, by dropping the 10% smallest towns, as measured by their pre-1990 employment. As shown in Table 3, none of these three changes qualitatively affect our baseline results. The only notable difference is that in the samples without Vienna and without the smallest towns, the total wage-bill effect is estimated to be significantly stronger in small towns than in large towns.

We also experiment with the definition of $Size_j$. In the fourth robustness test reported in Table 3, we replace the baseline definition by an inversely distance-weighted measure of a town's own employment

[10] See Nagy (2015) for a model of border changes affecting the spatial economy not only through the "local" market access channel we study in this paper but also through a "global" channel determined by changing relative distances to a country's dominant city.

Table 2
Baseline estimates
$Border_j$ defined by road distance from nearest CEEC border post or town (Czech Republic, Slovakia, Hungary or Slovenia).

Dependent variable	$\Delta(W^*E)$	ΔW	ΔE	$\Delta(W^*E)$	ΔW	ΔE	$\Delta(W^*E)$	ΔW	ΔE	$\Delta(W^*E)$	ΔW	ΔE
Definition of treatment sample (Border = 1)	≤ 25 km to nearest border post			≤ 35 km to nearest border post			≤ 60 km to nearest town ≥ 50,000			≤ 60 km to nearest town ≥ 20,000		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
$Fall_t \times Border_j \times Size_j$	−0.192	0.206**	−0.388***	−0.163	0.238**	−0.390***	−0.157	0.133*	−0.291***	−0.137	0.170***	−0.306***
	(0.117)	(0.094)	(0.098)	(0.146)	(0.117)	(0.114)	(0.121)	(0.063)	(0.110)	(0.124)	(0.065)	(0.106)
$Fall_t \times Border_i$	0.962***	0.207*	0.783***	1.005***	0.195*	0.859*	0.646*	0.115	0.567*	0.687**	0.024	0.691**
	(0.312)	(0.124)	(0.283)	(0.286)	(0.107)	(0.253)	(0.347)	(0.140)	(0.308)	(0.309)	(0.124)	(0.273)
$Fall_t \times Size_j$	0.013*	0.014**	−0.003	0.016**	0.014**	0.003	0.011*	0.015*	−0.002	0.012*	0.014**	−0.001
	(0.007)	(0.006)	(0.006)	(0.007)	(0.006)	(0.006)	(0.007)	(0.006)	(0.007)	(0.007)	(0.006)	(0.007)
No.towns for which $Border_i = 1$	317	317	317	522	522	522	325	325	325	393	393	393
R^2	0.149	0.056	0.181	0.149	0.056	0.181	0.149	0.056	0.181	0.149	0.056	0.181

Notes: quarterly data, 1976 Q1-2002 Q4; 2,047 towns, 221,076 observations; estimation with OLS; town and quarter fixed effects included throughout; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; *: $p = .1$, **: $p = .05$, ***: $p = .05$, ***: $p = .01$; "within" R^2, conditional on town fixed effects; W: normal wage, E: employment; $Size_j$: mean-differenced average pre-1990 town-level employment.

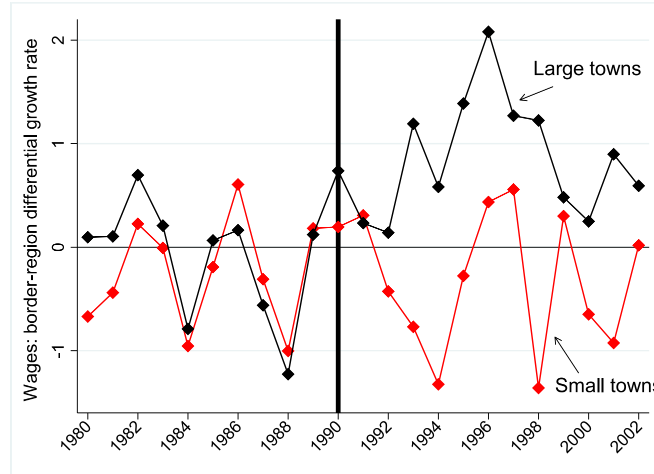

Fig. 5. Time profile of adjustment in wages large and small towns.
Notes: The graph shows the evolution from 1980 to 2002 of our estimated annual treatment effects (i.e. difference in wage growth rate between border and interior towns) in percentage points. Large towns: > 1000 employees in 1989 (black line); Small towns: ≤ 70 employees in 1989 (red line). For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.

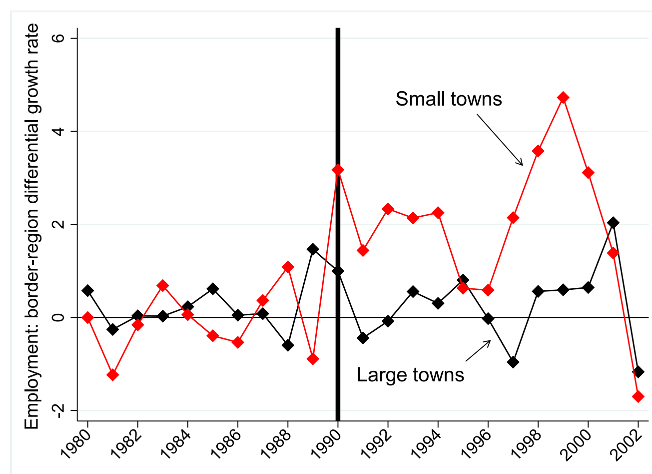

Fig. 6. Time profile of adjustment in employment large and small towns.
Notes: The graph shows the evolution from 1980 to 2002 of our estimated annual treatment effects (i.e. difference in employment growth rate between border and interior towns) in percentage points. Large towns: > 1,000 employees in 1989 (black line); Small towns: ≤ 70 employees in 1989 (red line). For interpretation of the references to color in this figure legend, the reader is referred to the web version of this article.

and that of its neighbours.[11] This measure is designed to take account of commuting and other spillover mechanisms among real-world towns (see Monte et al., 2016). In yet another definition, we compute $Size_j$ as employment density, dividing employment by constructible land area.[12] As shown in rows (D) and (E) of Table 3, our baseline results are robust to these variations in the definition of town size.

4.2.2. Mechanisms

In a series of further variants of our baseline model, we explore (and

[11] Specifically, we apply the standard centrality measure $Centrality_j = \sum_{m=1}^{M} E_m^{pre-1990} D_{jm}^{-2}$, where D_{jm} denotes the road distance between towns j and m, and $D_{jj} = 0.67\sqrt{area_j/\pi}$ (see e.g. Head and Mayer, 2010). The correlation between our benchmark measure of $Size_j$ and this centrality measure is 0.315.

[12] Constructible land area is defined as total area minus forest, water and uninhabitable mountain surfaces. The correlation between our benchmark measure of $Size_j$ and this density measure is 0.314. The correlation between the centrality and density measures is 0.991.

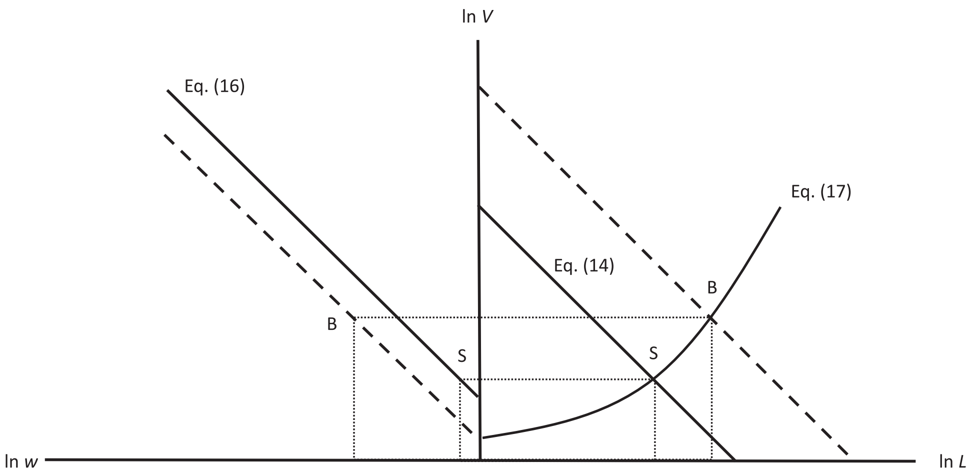

Fig. 7. Spatial equilibrium.
Notes: In trigonometric order, the first quadrant plots the common component of local (indirect) utility (vertical axis) as a function of the local population/workforce (horizontal axis), on log scales. Locations that command a higher utility level attract more workers (Eq. (17)). Competition for land implies that the relationship between V and L is negative (Eq. (14)): given population size, the more attractive location yields higher utility (dashed line). In equilibrium, the attractive location attracts a larger population/workforce (point 'B' vs. point 'S'). The second quadrant relates wages (horizontal axis) to the common level of utility (log scale). For a given utility level, the attractive location (dashed line) is associated with lower wages (compensating differential).

Table 3
Robustness: coefficients on triple interaction term ($Fall_t \times Border_j \times Size_j$).

Dependent variable:	$\Delta(W^*E)$ (1)	ΔW (2)	ΔE (3)	obs.
Repeat from Table 2:Baseline estimates ($Size_j$: mean-differenced average pre-1990 town-level employment)	−0.192	0.206**	−0.388***	221,076
	(0.117)	(0.094)	(0.098)	
(A) Baseline, controlling for Burgenland after 1995	−0.186	0.203**	−0.380***	221,076
	(0.118)	(0.093)	(0.098)	
(B) Baseline, dropping Vienna	−0.206*	0.161*	−0.363***	220,968
	(0.125)	(0.097)	(0.107)	
(C) Baseline, dropping 10% smallest towns	−0.245**	0.143**	−0.383***	196,668
	(0.104)	(0.060)	(0.090)	
(D) $Size_j$ defined as centrality (mean-differenced)	−4.555	4.275**	−8.347*	221,076
	(4.640)	(2.069)	(4.537)	
(E) $Size_j$ defined as density (mean-differenced)	−31.91	30.44**	−58.78*	221,076
	(33.70)	(14.69)	(33.53)	
(F) Baseline, controlling for unemployment	−0.164	0.163*	−0.317***	221,076
	(0.115)	(0.0864)	(0.093)	
(G) Baseline, long differences	−0.002	0.194***	−0.446***	4,062
	(0.006)	(0.051)	(0.116)	
(H) Baseline, long differences, blue-collar workers only	−0.004	0.113*	−0.798***	4002
	(0.007)	(0.059)	(0.295)	

Notes: quarterly data, 1976 Q1-2002 Q4; 2,047 towns in total; estimation with OLS; town and quarter fixed effects included throughout; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality clustering; *: $p = \ =0.1$, **: $p = 0.05$, ***: $p = 0.01$; $Border_j = 1$ if road distance from nearest CEEC border post ≤ 25 km (Czech Republic, Slovakia, Hungary or Slovenia); W: nominal wage, E: employment; see text for different definitions of $Size_j$ and specification of unemployment controls.

reject) three possible mechanisms behind the observed different responses across the town-size distribution: differential unemployment rates, skill compositions or sector-specific trends.

First, we augment the baseline models (1) to (3) by municipality-level pre-treatment *unemployment* rates and their interactions with $Fall_t$ and with $Fall_t \times Border_j$.[13] The implicit hypothesis is that the diagnosed more elastic local labor-supply schedules in smaller towns could be explained by higher unemployment rates in smaller towns. Such a mechanism based on initial unemployment rates, however, is doubly rejected by the data. First, the raw correlation between pre-treatment unemployment rates and $Size_j$ is in fact weakly positive: 0.04 for both our baseline definition of $Size_j$ and for density. Second, our estimated coefficients on the triple interaction, though slightly reduced in absolute size, retain their signs and statistical significance (Table 3, row F). Hence, different unemployment rates across different-sized towns do not seem to drive our results.

Another conceivable mechanism could be due to different skill compositions across different-sized towns and the associated comparative-advantage effects of trade liberalization. If Austria is skill abundant relative to its eastern neighbors, and large towns are skill abundant compared to small towns, then a standard Heckscher–Ohlin mechanism may give rise to the above-average wage increases in large towns. We therefore re-estimate our model for blue collar workers only.[14] As an additional robustness check, we show this specification estimated with total pre- and post-1990 growth rates instead of the quarterly series used so far. Row (G) shows that this change leaves the baseline estimates essentially unaffected. Finally, row (H) shows that above-average wage effects in large towns are found also when restricting the sample to blue collar workers. This effect is about a third smaller than in the baseline, suggesting that heterogeneous skill compositions may also play some role.

[13] We use town-level unemployment counts for 1971 and 1981, divided by town-level populations in those years. Being based on population censuses, these are the closest pre-treatment years for which town-level data are available.

[14] Austrian labor law distinguishes two types of employee contract, "Arbeiter" and "Angestellte", and this distinction is reported in the ASSD. The former contract is used exclusively for low-skill manual workers and thus allows us to restrict the sample to blue collar workers. The latter contract was historically reserved to white collar jobs but has become used more broadly. As a result, the residual category of workers not part of our blue collar category contains a mixture of blue and white collar workers and is therefore uninformative for an analysis by skill group.

Table 4
Matching: treatment and control towns matched on pre-1990 primary/secondary/tertiary employment shares.

Dependent variable:	$\Delta(W^*E)$ (1)	ΔW (2)	ΔE (3)
	All towns (2,001 obs.)		
$Fall_t \times Border_j$	0.899***	0.280***	1.184**
	(0.206)	(0.086)	(0.208)
	Large towns (> 1000 employees pre-1990; 174 obs.)		
$Fall_t \times Border_j$	1.378***	0.573***	1.305***
	(0.345)	(0.125)	(0.558)
	Small towns (= 70 employees pre-1990; 636 obs)		
$Fall_t \times Border_j$	1.237**	0.336	1.834***
	(0.440)	(0.206)	(0.698)

Notes: reported coefficients are average treatment effects on pre-1990 and post-1990 average annual growth rates; standard errors in parentheses; * : $p= .1$, **: $p= .05$, ***: $p= .01$; $Border_j = 1$ if road distance from nearest CEEC border post ≤ 25 km (Czech Republic, Slovakia, Hungary or Slovenia); W: nominal wage, E: employment.

Austria's eastern border regions were and remain economically less developed than most other Austrian regions. Hence, differential wage and employment trajectories between border and interior towns could be due to sector-specific trends rather than the impact of trade liberalization. Figs. 5 and 6 suggest that pre-1990 wage and employment growth in border towns did not systematically diverge from growth rates observed in interior towns. We address this issue more formally by estimating average treatment effects of $Fall_t \times Border_j$ after matching each border town with up to two interior towns that resemble the border town most closely in terms of their pre-treatment employment distributions across primary, secondary and tertiary activities.[15] In the first panel of Table 4, we show that the matching procedure does not undo the detected treatment effects on average border-town wages and employment. In the second and third panels of Table 4, we compute treatment effects separately for large towns and for small towns. It becomes apparent again that the wage effect is stronger in large towns while the employment effect is somewhat larger in small towns. In this instance, however, the wage-bill effect is essentially identical in the two sub-samples.

4.2.3. Placebo test

By way of an alternative explanation for our central findings, one might suspect that in an era of expanding cross-border trade and rapid European integration, border regions generally fared better than interior regions, and that the effects we attribute to the opening of central and eastern European economies in fact were generic features of border regions in the post-1990 period. We examine this proposition by re-estimating our baseline empirical model augmented by a placebo treatment group, defined as towns within 25 km from the nearest road border crossing with one of Austria's western neighbor countries, Germany, Italy, Switzerland or Liechtenstein. We add this placebo treatment to the baseline specification, and we introduce it on its own, omitting the eastern border towns from the sample. The estimation results are presented in Table 5. Controlling for the placebo group does not qualitatively affect our estimates for the original treatment, and we find no statistically significant coefficients on the triple interaction term in the placebo treatment. We even observe that western border regions experienced significantly below-average employment growth post-1990. This clearly shows that the post-1990 gains in eastern border regions did not reflect a positive employment trend in border regions in general. Hence, our placebo results support the case for interpreting our baseline findings as the causal effects of trade liberalization induced by the fall of the Iron Curtain.

5. A multi-town model of spatial adjustment

In this section we develop a model that puts forth a plausible mechanism, consistent with our empirical findings. This model will enable us to conduct some welfare analysis in Section 6. Specifically, we model the spatial economy using three building blocks, similarly to Redding (2016): a Helpman–Krugman economic geography *à la* Redding and Turner (2015), using mostly their notation; production amenities as in

Table 5
Placebo tests
$WestBorder_j$ defined as within 25 km road distance from nearest border point with Italy, Switzerland, Liechtenstein or Germany
$Border_j$ defined as within 25 km road distance from nearest border post with Czech Republic, Slovakia or Hungary.

Dependent variable:	$\Delta(W^*E)$	ΔW	ΔE	$\Delta(W^*E)$	ΔW	ΔE
Sample:	Full sample			Without Eastern border towns		
	(1)	(2)	(3)	(4)	(5)	(6)
$Fall_t \times Border_j \times Size_j$	−0.188	0.206**	−0.385***			
	(0.117)	(0.094)	(0.098)			
$Fall_t \times WestBorder_j \times Size_j$	0.259	0.071	0.199	0.258	0.0745	0.194
	(0.203)	(0.111)	(0.140)	(0.203)	(0.112)	(0.139)
$Fall_t \times Border_j$	0.809**	0.215*	0.622**			
	(0.320)	(0.126)	(0.290)			
$Fall_t \times WestBorder_j$	−0.907***	0.052	−0.958***	−0.907***	0.041	−0.948***
	(0.300)	(0.118)	(0.267)	(0.302)	(0.118)	(0.269)
$Fall_t \times Size_j$	0.008	0.014**	−0.005	0.008	0.014**	−0.005
	(0.007)	(0.006)	(0.008)	(0.007)	(0.006)	(0.008)
No. towns for which $WestBorder_j = 1$	301	301	301	301	301	301
No. towns for which $Border_j = 1$	317	317	317	0	0	0
R^2	0.149	0.056	0.181	0.149	0.056	0.181

Notes: quarterly data, 1976 Q1 - 2002 Q4; 2047 towns, 221,076 observations; estimation with OLS; town and quarter fixed effects included throughout; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; *: $p= .1$, **: $p= .05$, ***: $p= .01$; "within" R^2, conditional on town fixed effects; W: nominal wage, E: employment; $Size_j$: mean-differenced average pre-1990 town-level employment.

[15] The ASSD data also contain information on somewhat more disaggregated sector affiliations, but we found this information to be too noisy to be informative.

e.g. Roback (1982); and heterogeneous preferences over locations and discrete location choices following Luce (1959) and Ben-Akiva and Lerman (1985) as in e.g. Behrens et al. (2017).

5.1. Model

We consider a population $\overline{L}$ of individuals supplying one unit of labor each and choosing to live and work among a continuum of Austrian 'towns' that are heterogeneous in their land supply and production amenities, as well as in their location in the trade-cum-transportation network. Austrian towns trade with each other and with Foreign locations. There is no international migration. Land is used in production and in housing services. We denote by $\mathscr{I}$ the set of all locations and by $[0, 1] \subset \mathscr{I}$ the set of Austrian locations. We sometimes denote by i the town in which production takes place and by n the town in which consumption takes place.

5.2. Preferences and technology

Individual preferences are defined over local natural amenities, local (non-traded) housing services H, and a Dixit–Stiglitz consumption basket of tradable goods C, where the elasticity of substitution among varieties of the composite good C is constant and denoted by $\sigma > 1$. Individuals have idiosyncratic valuations of local amenities denoted by ϵ; by contrast, preferences over C and H are common to all. We can thus write:

$$\widetilde{U}(C, H, \epsilon) = U(C, H)\epsilon, \quad \text{where} \quad U(C, H) \equiv \mu \ln C + (1-\mu)\ln H \tag{4}$$

is the common component of utility and $\mu \in (0, 1)$ is the expenditure share on good C.

We assume that the idiosyncratic terms $\ln \epsilon$ are distributed iid Gumbel with standard deviation $\pi/\sqrt{6}$, as in Luce (1959). As a result, the probability that a randomly drawn individual chooses to locate in town $n \in [0, 1]$ follows a continuous logit, as in Ben-Akiva and Lerman (1985):

$$\mathbb{P}\mathrm{r}(n) = \frac{U(C(n), H(n))}{\int_0^1 U(C(i), H(i))\mathrm{d}i}, \tag{5}$$

where $C(n)$ and $H(n)$ denote utility maximizing consumption levels in town n.

Production of the Dixit–Stiglitz composite good features increasing returns to scale at the plant level and requires labor L and 'production structures' K. Specifically, we write the production function of the representative firm in town i as

$$\ell(i)^{\alpha} k(i)^{1-\alpha} = F + \frac{x(i)}{A(i)}, \tag{6}$$

where ℓ is the workforce of the representative firm, k is the size of its structure, x its output, and $F > 0$ and $\alpha \in (0, 1)$ are parameters. A, the marginal product of the composite input, is a location-specific production amenity. Sector C is monopolistically competitive. Consequently, firms charge a constant markup $\sigma/(\sigma - 1)$ over marginal costs, and free entry and exit of firms drive profits to zero. The firm size consistent with zero profit is $x(i) = A(i)F(\sigma - 1)$.

Both housing and production structures are produced using land, T. We assume that housing can be converted into production structures and vice-versa with a constant elasticity of transformation $\tau > 1$:

$$\overline{T}(n)^{1-1/\tau} = H(n)^{1-1/\tau} + K(n)^{1-1/\tau}. \tag{7}$$

The stock of land, $\overline{T}(n)$, is exogenously given.[16] Maximizing land value $r(n)H(n) + s(n)K(n)$ under the constraint (7) implies that H and K are proportional to $\overline{T}$ in equilibrium,

$$K(n) = \overline{T}(n)\left[\frac{\mu(1-\alpha)}{1-\alpha\mu}\right]^{\tau/(\tau-1)}, \qquad H(n) = \overline{T}(n)\left[\frac{1-\mu}{1-\alpha\mu}\right]^{\tau/(\tau-1)} \tag{8}$$

by (10) below; the factor prices r and s absorb other locational differences.

Let $v(n)$ denote per-capita nominal earnings in n so that

$$v(n)L(n) = w(n)L(n) + r(n)H(n) + s(n)K(n), \tag{9}$$

where w, r, and s denote the unit prices of labor, housing, and production structures, respectively. Using the Cobb–Douglas properties of preferences and production in equations (4) and (6) yields

$$v(n) = \frac{w(n)}{\alpha\mu} \quad \text{and} \quad \frac{w(n)L(n)}{\alpha} = \frac{s(n)K(n)}{1-\alpha}. \tag{10}$$

5.3. Trade and market access

Trade from i to n is costly and parameterized by the matrix of iceberg trade costs $d(n, i) > 1$ for all $n \neq i$ and $d(n, n) = 1$ for all n.

Following Redding and Venables (2004); Redding and Turner (2015), we can show that firms located in i break even if and only if the unit cost of production obeys

$$w(i)^{\alpha} s(i)^{1-\alpha} = \kappa_0 A(i)^{1-1/\sigma} fma(i)^{1/\sigma}, \tag{11}$$

where $\kappa_0 > 0$ collects parameters and fma denotes 'firm market access' and is defined as

$$fma(i) \equiv \int_{n\in\mathscr{I}} d(n, i)^{1-\sigma} v(n)L(n)P(n)^{\sigma-1}\mathrm{d}n, \tag{12}$$

and where P is the CES price index of the composite Dixit–Stiglitz good,

$$P(n) \equiv \left[\int_{i\in\mathscr{I}} M(i)d(n, i)^{1-\sigma}p(i)^{1-\sigma}\mathrm{d}i\right]^{1/(1-\sigma)},$$

$p(i)$ is the fob price charged by firms located in i, and $M(i)$ is the equilibrium mass of varieties produced in i. It follows by inspection of (11) that towns endowed with desirable production amenities and high firm market access pay higher wages and/or production structure prices. By the same token, consumers located in a town with good access to production centers, i.e. living in towns with a low $d(n, \cdot)$ on average, face lower consumer prices than consumers living in remote towns. Following Redding and Turner (2015), we thus define 'consumer market access' as

$$cma(n) \equiv P(n)^{1-\sigma} = \int_{i\in\mathscr{I}} M(i)d(n, i)^{1-\sigma}p(i)^{1-\sigma}\mathrm{d}i. \tag{13}$$

The deterministic component of indirect utility, i.e. the dual of U in (4), is

$$\ln V(n) = \ln v(n) - \mu \ln P(n) - (1-\mu)\ln r(n).$$

Using the definition for cma in (13), the equilibrium relationship between factor rewards and fma in (12), the equilibrium relationship among factor rewards and per-capita earnings in (10), and the full-employment condition for land in (7), we can rewrite this expression as

$$\ln V(n) = \kappa_L + \frac{\sigma-1}{\sigma}\mu \ln A(n) + \mu \ln ma(n) - (1-\alpha\mu)\ln\frac{L(n)}{\overline{T}(n)}, \tag{14}$$

where κ_L collects parameters and $ma(n)$ is overall market access,

[16] Two comments are in order. First, the formulation in (7) encompasses two standard alternative classes of models, Roback (1982) and Redding and Turner (2015). Land used for housing development and land used for industrial development are perfect substitutes in Roback (1982), so that $\tau \to \infty$. In Redding and Turner (2015), only firms use land, so that $\alpha = 1$, which implies $K(n) = 0$ and, in turn, $H(n)$ is exogenously given. Second, we can easily relax the assumption that land supply is inelastic: all results go through if land supply is iso-elastic in land prices.

$$\ln ma(n) \equiv \frac{1}{\sigma}\ln fma(n) + \frac{1}{\sigma-1}\ln cma(n). \tag{15}$$

Equation (14) subsumes important mechanisms at work in the model. First, towns endowed with a high exogenous productivity are attractive because they can pay higher wages. Second, towns with good market access offer high consumer utility because they are able to pay high wages or reduce the cost of consuming the composite good, or both. Finally, densely populated places incur congestion, which raises land prices and thereby hurts the wellbeing of residents. We turn to this mechanism next.

Equation (14) is akin to a **local labor demand** function in the (L, V)-space: the marginal contribution of an additional worker/resident is decreasing because of decreasing returns to labor in the production of the differentiated good and crowding on the housing and land markets.

The first quadrant of Fig. 7 illustrates (14), which provides a negative relationship between town size L and the town-specific common component V of utility, given market access. An increase in ma, A, or $\overline{T}$ all shift this schedule outwards. Of course, L and ma are endogenous variables, which means that we need another set of conditions to complete the characterization of the equilibrium, which we provide in Section 5.4 below.

One advantage of the data we exploit in our empirical work is that we observe town-specific median wages. It is therefore worth characterizing the theoretical relationship between wages and market access. In particular, we can use (8), (10), (11), (13), and (14) to obtain an equilibrium relationship between w and V:

$$\begin{aligned}\ln w(n) &= \kappa_w + \frac{1-\alpha}{1-\alpha\mu}\ln V(n) + \frac{(\sigma-1)(1-\mu)}{\sigma(1-\alpha\mu)}\ln A(n) \\ &\quad + \frac{1-\mu}{\sigma(1-\alpha\mu)}\ln fma(n) - \frac{(1-\alpha)\mu}{(\sigma-1)(1-\alpha\mu)}\ln cma(n),\end{aligned} \tag{16}$$

where κ_w collects parameters. In words, utility and wages are positively correlated, and wages are increasing in local production amenities. Both properties are in line with economic intuition. An additional property of (16) is particularly noteworthy: market access has a theoretically ambiguous effect on wages. On the one hand, *firm* market access has a positive effect on wages because a high fma enables firms to pay high wages and yet break even. On the other hand, *for a given utility level V*, a better *consumer* market access is negatively capitalized into wages by the logic of compensating differential popularized by e.g. Roback (1982). The former effect is likely to dominate the latter if the expenditure share of tradable goods C is low relative to the share of non-tradable goods H, i.e. if μ is low enough. The second quadrant of Fig. 7 plots (16) for the same arbitrary pair of towns as in the first quadrant. Our graph illustrates the case in which the productivity and firm market access advantages of the large town dominate its consumer market access disadvantage, as is the case in the data.

5.4. Equilibrium

Labor-market equilibrium requires the *actual* number of workers/residents in n to be equal to the number of workers/residents *wishing* to live there:

$$\forall\, n \in [0,1]: \quad L(n) = \mathbb{P}\mathrm{r}(n)\overline{L} = \frac{U(n)}{\int_0^1 U(i)\mathrm{d}i}\overline{L} = \frac{\ln V(n)}{\ln \mathbb{V}}\overline{L}, \tag{17}$$

where $\mathbb{V}$ is the expected utility of a utility-maximizing Austrian resident, which is equal to the geometric average of the deterministic component of utility across Austrian towns:

$$\ln \mathbb{V} = \int_0^1 \ln V(i)\mathrm{d}i.$$

Equation (17) yields a positive relationship between (indirect) utility $V(n)$ and population $L(n)$. It is akin to a **local labor supply** in (L, V)-space: a higher real income V attracts more workers/residents. It then follows that the elasticity of labor supply is decreasing in town size L,

$$\begin{aligned}\eta(n) &\equiv \left.\frac{\partial \ln L(n)}{\partial \ln V(n)}\right|_{\text{Eq. (17)}} \\ &= \frac{\overline{L}}{\ln \mathbb{V}}\frac{1}{L(n)},\end{aligned} \tag{18}$$

where the second equality follows from (17). The labor supply of large towns is less elastic than that of small towns because the marginal utility of income is decreasing due to the concavity of U in (4), which implies that the valuation of natural amenities increases with (real) income.

A unique general equilibrium with a stable, non-degenerate distribution of population across towns exists under the assumption that bilateral transportation costs are symmetric and if dispersion forces arising from land scarcity dominate agglomeration forces due to returns to scale at the firm level and costly trade.[17] The first quadrant of Fig. 7 illustrates the equilibrium for two arbitrary towns. The upward sloping local labor supply curve plots equation (17) and is the same for both towns. The downward sloping local labor demand schedules plot equation (14) for a town endowed with relatively low values of A, ma, or $\overline{T}$ (the solid line), and for a town endowed with relatively high levels of these variables (the dotted line). The schedules intersect at points S (for 'small') and B (for 'big'), respectively. In equilibrium, then, the town enjoying relatively poor market access, low production amenities, and small land endowment is smaller than the town endowed with more of any of these. Importantly, the upward sloping curve is convex so that it is steeper at point B than at point S, a property that plays an important role below.

5.5. Interpreting the empirical results

In characterizing the equilibrium, we treat fma and cma as right-hand side variables, which allows us to express changes in endogenous variables of interest, such as wage rates and the number of workers, as functions of the shocks in market access brought about by the fall of the Iron Curtain, as we do in our empirical work (where in addition we control for initial conditions and town-specific trends).

Let 'hats' denote log-changes and subscripts '0' denote initial values in levels. We think of the fall of the Iron Curtain as $\widehat{fma}(n) \geq 0$ and $\widehat{cma}(n) \geq 0$ for all $n \in [0, 1]$. This difference over time is the first 'diff' in our triple difference empirical strategy.

The second 'diff' considers two towns that are initially identical in all respects but one: town b is a 'treatment town' located near the border, while town c is a 'control town' located in the interior. Let $\Delta\hat{x} \equiv \hat{x}(b) - \hat{x}(c)$ denote the difference in the change of any variable x between the border town and the interior town. In particular, we assume $\Delta\widehat{fma} > 0$ and $\Delta\widehat{cma} > 0$.[18] Totally differentiating (14), (16), and (17), we obtain the following expression for the relative effect of market access on town size:

$$\Delta\hat{L} = \frac{\mu}{\eta_0^{-1} + (1-\alpha\mu)}\Delta\widehat{ma}, \tag{19}$$

[17] The method of proof follows Allen and Arkolakis (2014), Redding and Turner (2015) and Redding (2016). The assumption that bilateral transport costs are symmetric is sufficient but not necessary; see Redding (2016). In our model, the commonly assumed 'no black hole condition' (e.g. Helpman, 1998) holds if $\sigma(1-\alpha\mu) > 1$, which we henceforth assume so that the model features a unique, stable equilibrium. Agglomeration forces are decreasing in σ because individual firms are large (and unexploited scale economies low) when varieties are close substitutes; dispersion forces are strong when the shares of land in production and consumption are high ($1-\alpha$ and $1-\mu$, respectively).

[18] Of course, all towns are treated in a general equilibrium because all towns in the network are linked by trade and internal migration (Redding and Turner, 2015). We only identify a relative treatment effect: border towns are 'more treated' than interior towns, which we capture by our assumption $\Delta\widehat{fma} > 0$ and $\Delta\widehat{cma} > 0$.

where $\Delta\widehat{ma} \equiv \sigma^{-1}\Delta\widehat{fma} + (\sigma-1)^{-1}\Delta\widehat{cma} > 0$ by (15). Thus, under the assumption that border towns get a larger market access shock than interior towns, the labor force of the former grows relative to the labor force of the latter, as we find in our main results reported in Section 4.1.

By the same token, we find:

$$\Delta\hat{w} = \left[1 - \frac{(1-\alpha)\mu}{\eta_0^{-1} + (1-\alpha\mu)}\right]\frac{\Delta\widehat{fma}}{\sigma} - \frac{(1-\alpha)\mu}{\eta_0^{-1}+(1-\alpha\mu)}\frac{\Delta\widehat{cma}}{\sigma-1}. \tag{20}$$

The two sources of market access changes have opposite effects on wages, as explained above. Empirically, we find $\Delta\hat{w} > 0$. We interpret this result as evidence that the effect of an improvement in firm market access dominates the effect of an improvement in consumer market access.

The third and final 'diff' of our triple difference empirical strategy involves comparing the effects of the fall of the Iron Curtain on large versus small towns. To this aim, observe that the coefficients of the $\Delta\widehat{ma}$'s in (19) and (20) all depend on the pair-specific labor supply elasticity, η_0, which is decreasing in initial town size by (18). Let L_0 denote the pre-shock size of an arbitrary town. Using (18), we obtain the following cross derivatives by inspection of (19) and (20), respectively:

$$\frac{\partial^2\Delta\hat{L}}{\partial\widehat{ma}\partial L_0} < 0, \qquad \frac{\partial^2\Delta\hat{w}}{\partial\widehat{fma}\partial L_0} > 0, \quad \text{and} \quad \frac{\partial^2\Delta\hat{w}}{\partial\widehat{cma}\partial L_0} > 0. \tag{21}$$

That is to say, following the market access shock, wages are expected to absorb a larger fraction of the shock in large towns than in small towns; by contrast, employment adjusts proportionally more strongly in small towns than in large towns. Analogously, we may also consider how the wage bill at the town level, $w(n)L(n)$, evolves as a result of changes in market access. Using (19) and (20), and the definition of $\Delta\widehat{ma}$ yields

$$\Delta\widehat{wL} \equiv \Delta\hat{w} + \Delta\hat{L} = \frac{\Delta\widehat{fma}}{\sigma} + \frac{\alpha\mu}{\eta_0^{-1} + (1-\alpha\mu)}\Delta\widehat{ma},$$

so that

$$\frac{\partial^2\Delta\widehat{wL}}{\partial\widehat{ma}\partial L_0} < 0. \tag{22}$$

Hence, the sum of the price and quantity response to a market access shock is expected to be decreasing in initial town size. The heterogeneous effects in Eqs. (21) and (22) are weakly supported by our central findings of Section 4.1.

Figure 8 illustrates the effect of a positive market access shock on the big town, B, and small town, S, of Fig. 7. The axes of Fig. 8 report the *first differences* of the variables on the corresponding axes of Fig. 7. For instance, the downward sloping schedule in the first quadrant corresponds to (14) in first differences, i.e. $\hat{V} = -(1-\alpha\mu)\hat{L}$. This schedule is identical for both B and S. The positive border shock shifts this schedule in parallel fashion, allowing for a larger employment-cum-population size and/or a higher utility: the mathematical definition of the dashed downward sloping schedule is $\hat{V} = \mu\widehat{ma} - (1-\alpha\mu)\hat{L}$ (recall that we assume $\widehat{ma}$ to be larger for border towns than for interior towns). The upward sloping schedules in the first quadrant of Fig. 8 plot the local labor supply, i.e. the slope of the upward sloping curve of Fig. 7. As explained above, they are proportional to the pre-shock size L_0 by (18). Hence, an identical market access shock has a stronger employment effect in the initially small town than on the initially large town. Conversely, the wage effect is larger for the large town than for the small town.

To summarize, the model makes the following theoretical predictions that are consistent with the data:

1. Following a market access shock such as the fall of the Iron Curtain, border towns experience an increase in employment and wages relative to interior towns (see Eq. (19) and (20) above).
2. The coefficient of the interaction between treatment and town size is negative when the dependent variable is the change in town employment (see Eq. (21) above).
3. The coefficient of the interaction between treatment and town size is positive when the dependent variable is the change in the town

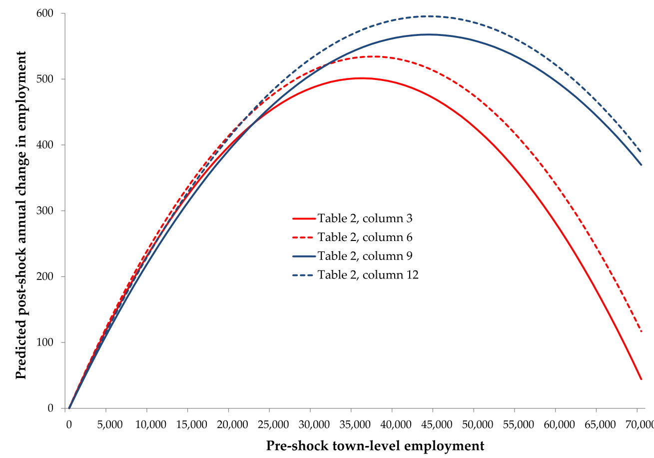

Fig. 9. Predicted trade-induced absolute changes in town-level employment.
Notes: Predicted absolute annual employment changes in border towns post-1990 as implied by the coefficients of our four baseline estimations reported in Table 2 (constant term = 2.00; mean pre-1990 border-town employment = 502). For example, for a town of the size of Villach, with around 14,000 employees in 1989, the graph predicts a cumulative post-shock differential (relative to an interior town) increase in employment of 300.

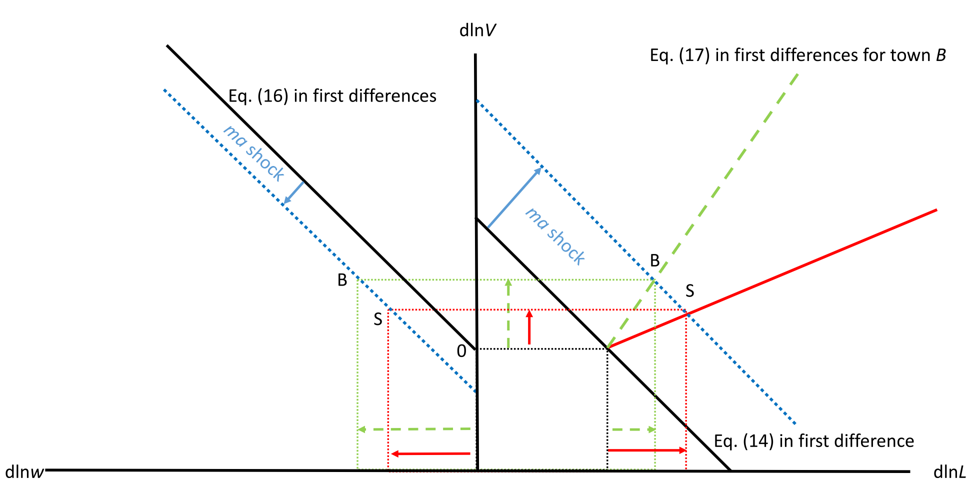

Fig. 8. Adjustment to a market access shock.
Notes: In trigonometric order, the first quadrant plots the variation in the common component of local (indirect) utility (vertical axis) as a function of the variation in local population/workforce (horizontal axis), on log scale s, following a positive shock to market access. Locations that are initially small have a more elastic labor supply than large locations and grow relatively more following a shock to market access. The second quadrant relates changes in wages (horizontal axis) to changes of the common utility level (log scale). For a given change in utility, the wage rate in the attractive location (dashed line) grows relative to the wage rate in the small, less attractive location.

wage (see Eq. (21) above).

4. The coefficient of the interaction between treatment and town size is negative when the dependent variable is the change in the town wage bill (see Eq. (22) above).

The first prediction is consistent with the empirical findings of Redding and Sturm (2008) and Brülhart et al. (2012). The other three are novel and underpin our triple-difference estimation strategy.

6. Welfare

Our empirical results support the central qualitative prediction of the model: large towns respond primarily through wage changes while small towns respond primarily through employment changes. Insofar as it validates our theoretical model, this positive result also implies normative predictions regarding welfare changes across different-sized towns.

In the model, the relative change of the common component of utility in a given town is proportional to the absolute change in that town's employment level.[19]

This result can be seen by totally differentiating (14), (16), and (17), which yields:

$$\Delta\widehat{V} = \eta_0^{-1}\Delta\widehat{L} = \frac{\ln \mathbb{V}_0}{\overline{L}}\Delta \mathrm{d}L, \tag{23}$$

where the second equality follows from (18), and $\Delta \mathrm{d}L \equiv \mathrm{d}L(b) - \mathrm{d}L(c)$ is the change in the population of the border town relative to the control town. That is to say, the relative change of the common component of utility is proportional to the absolute change in employment.[20]

As a town grows in size, its marginal residents are increasingly reluctant to settle there and need a commensurately increasing monetary compensation. This is apparent in Fig. 8. Importantly, to the extent that the *absolute* change in employment tends to be larger in large towns than in small towns in the data, a given market-access shock will benefit residents of large towns more than it benefits those of small towns. This is due to the distribution of the idiosyncratic component of utility for individual locations, which implies that the elasticity of the local labour supply (migration) is decreasing in town size.

Our regression specifications consider relative changes. In absolute terms, bigger towns may well experience larger employment changes. We therefore use the coefficient estimates of our four baseline employment regressions of Table 2 to compute the implied absolute employment changes along the interval of observed town sizes. These predicted values are shown in Fig. 9. Our results suggest a non-monotonic relationship within the range of observed town sizes: the largest differential border-town gains to local employment – and thus welfare – induced by the trade shock are predicted for border towns with observed employment of roughly 40,000, averaged over the 1976–1989 period, which corresponds to a population of around 150,000.[21]

The observed decrease in absolute employment gains of towns above a certain size is consistent with heterogeneous treatment intensity across the town-size distribution: at least above a certain size threshold, larger towns trade relatively more with themselves and are therefore relatively less affected by a given shock to external market access.[22]

7. Conclusions

We have explored the impact of trade liberalization on employment and wage growth of different-sized towns using a quasi-experimental setting provided by Austrian regions. The fall of the Iron Curtain in 1990 represents a large and fully exogenous trade shock for the Austrian economy. Eastern Austrian towns being more exposed to this shock than interior towns, we have defined eastern border regions as the treatment group and the rest of Austria as the control group. We detect significant heterogeneity in treatment effects across the size distribution of towns. Larger towns are found to have larger nominal wage responses, but smaller employment responses, than smaller towns.[23]

This pattern of adjustment turns out to be consistent with a multi-region model, in which responses to external trade liberalization take place via internal labor migration among heterogeneous towns. This model implies that local labor supply schedules are more elastic in small towns than in large towns and therefore predicts that trade liberalization will trigger stronger wage effects in large towns and stronger employment effects in small towns. The model also directly leads to our triple-difference empirical strategy.

The positive predictions of our model are borne out by the experience of Austrian towns. That same model implies a positive relationship between a town's absolute employment change and the welfare change for its incumbent residents. Our estimates imply a non-monotonic relationship between town size and the trade-induced increase in employment levels, with a peak for towns with a population of around 150,000. This suggests that residents of medium-sized towns gain the most from a given opening of cross-border trade.

An important question remains: to what extent do the increases in employment and wages experienced by border towns come at the expense of other towns (in the interior of the same country, or across the border), and to what extent do they reflect net increases in economic activity? Our difference-in-difference empirical appraoch does not shed light on this question, and our theoretical model assumes a fixed level of employment. The net effects of trade-induced differential regional growth patterns therefore remain an open issue for future research.

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[19] In comparing two equilibria, changes in the welfare of stayers are identical to changes to the common component of utility, which can differ across town types. Overall, however, the expected utility of all individuals in the economy is equal to the expected utility conditional on living in any town n, so that changes in $\mathbb{V}$ fully capture changes in the expected utility of stayers and movers combined (this knife-edge result is a consequence of the assumption that idiosyncratic location preferences take an extreme value distribution; see e.g. Redding, 2016).

[20] Two properties of (23) are noteworthy. First, as a town grows in size, its marginal residents are increasingly reluctant to settle there and they need an equally increasing monetary compensation. Thus, $\widehat{V}$ is an upper bound of the per capita increase in welfare in towns that gain population (and a lower bound of the welfare loss in towns that lose population). Second, insofar as dL is larger in large towns than in small towns, the market-access shock benefits large towns more than it benefits small towns – with the caveat just raised when using metric V.

[21] Total employment recorded in our data corresponds to some 26 to 27 percent of Austrian population over our sample period.

[22] The fact that we find big treatment effects for Vienna, by far Austria's largest "town", may in turn be due to the fact that Vienna was particularly well connected to the eastern markets via transport, communication and organizational networks.

[23] We have in addition found evidence of trade liberalization boosting housing prices in border towns (see Brülhart et al., 2015).

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