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How wages and employment adjust to trade liberalization: Quasi-experimental evidence from Austria ☆

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ABSTRACT

We study the response of regional employment and nominal wages to trade liberalization, exploiting the natural experiment provided by the opening of Central and Eastern European markets after the fall of the Iron Curtain in 1990. Using data for Austrian municipalities, we examine differential pre- and post-1990 wage and employment growth rates between regions bordering the formerly communist economies and interior regions. If the 'border regions' are defined narrowly, within a band of less than 50 km, we can identify statistically significant liberalization effects on both employment and wages. While wages responded earlier than employment, the employment effect over the entire adjustment period is estimated to be around three times as large as the wage effect. The implied slope of the regional labor supply curve can be replicated in an economic geography model that features obstacles to labor migration due to immobile housing and to heterogeneous locational preferences.

1. Introduction

We address a fundamental but to date surprisingly underresearched question: how do changes in market access affect factor prices and factor quantities? To put it simply: if a certain region offers advantageous access to markets elsewhere, will this advantage translate into a large number of producers locating in that region, will it translate into higher factor rewards for producers located there, or will we observe some of both effects? As a natural corollary to this question, we also study such effects across different time horizons, as quantity and price adjustments may well materialize at different speeds. We focus on the case where changes in market access are due to the liberalization of international trade.

Why should we care about the difference between factor price effects and factor quantity effects of changes in market access? First, this distinction helps us understand adjustment mechanisms of regional economies, by allowing us to trace regional factor supply schedules. For example, large price effects suggest the existence of important barriers to the reallocation of labor and capital across space and/or across sectors. Information on the relative magnitude of price and quantity effects can thereby help us gauge the realism of alternative theoretical models. Second, the policy implications of market-access effects vary considerably depending on whether these effects work through factor prices or through factor quantities. Price effects bring about spatial inequality of (pre-tax) factor rewards, which can potentially be evened out via redistributive policy. Quantity effects may imply problems from congestion in central locations and depopulation in peripheral ones, or from specialization patterns that make regions vulnerable to sector-specific shocks.

Almost all research to date has focused on the two polar cases, by looking either at quantity effects or at price effects, thus implicitly assuming regional factor supply schedules to be either horizontal or vertical. Many empirical studies that are formally linked to the theory assume that intersectoral and/or interregional factor supplies are infinitely elastic, which leaves room for quantity effects only. The

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sizeable empirical literature on home-market effects, initiated by Davis and Weinstein (1999), belongs to this category. Redding and Sturm (2008) were first to explore quantity adjustment using a natural experiment involving changes in market access, by tracking changing populations of cities located along the border between East and West Germany during the country's division and reunification in the 20th century. Faber (2009) has studied the effects of highway construction in China on industrial production of rural counties to identify the causal effect of market access on regional output. Conversely, a strand of the literature due mainly to Hanson (1997, 2005) has assumed that factor supplies are inelastic, such that market-access effects manifest themselves solely via factor prices (i.e. wages). Redding and Venables (2004) have used this approach to study the determinants of international differences in per-capita income and found that the geography of access to markets and sources of supply is quantitatively important.

When specifically studying intra-national adjustment to international trade liberalization, most researchers have looked at quantity effects, mainly in terms of city populations (e.g. Ades and Glaeser, 1995; Henderson, 2003) and of regional employment (e.g. Hanson, 1998; Brülhart et al., 2004; Sanguinetti and Volpe Martincus, 2009). A smaller number of researchers have alternatively considered price effects, in terms of regional wages (e.g. Hanson, 1997; Chiquiar, 2008). The combination of quantity and price effects has not yet, to our knowledge, been studied in this context.[1]

The theoretical distinction between price and quantity effects of market access has been brought into focus by Head and Mayer (2004). Using a geography model featuring imperfectly elastic factor supply to the sector that is subject to agglomeration forces, they showed that, depending on the size of this elasticity, quantity effects or price effects may dominate. In a subsequent paper (Head and Mayer, 2006), they have investigated this issue empirically, by estimating how European region-sector wages deviate from a benchmark pattern that would be consistent with pure quantity responses to agglomeration forces. They found stronger evidence for price effects than for quantity effects. They acknowledged that, while their strategy for estimating wage responses was fully structural, the estimation of employment changes had to rely on *ad hoc* regressions, and that their empirical implementation faced considerable challenges in terms of measurement and causal inference.

Our approach is to draw on a natural experiment and to use a difference-in-difference identification strategy. We take the fall of the Iron Curtain in 1990 as an exogenous event that increased overall market access of Austrian regions, but more so for regions close to Austria's eastern border. Comparing post-1990 wage and employment growth in border regions to that in interior regions, we can control for common shocks and isolate the effects of increased market access with considerable confidence. This quasi-experimental strategy obviates the need to construct an artificial benchmark that would have to be tied to a specific variant of the underlying model and would inevitably be prone to measurement error.

Our central contribution is to consider factor-price effects as well as factor-quantity effects. Specifically, we trace the impact of improved market access on both nominal wages and employment levels. We find that the employment effect exceeds the wage effect by a factor of around three. Furthermore, we are able to characterize the time profile of adjustment along those two margins, observing that wage rises precede the increases in employment.

In addition, we seek to replicate our estimated ratio of employment-to-wage-adjustment in a calibrated three-region economic geography model. A non-tradeable housing sector acts as a dispersion force against the agglomeration tendencies that arise from the interplay of trade costs, product differentiation and increasing returns. When we add a further dispersion force due to heterogeneous locational preferences, we find that the model predicts our central estimate of relative labor-market adjustment margins for realistic parameter values.

The remainder of the paper is organized as follows. In Section 2, we present a theoretical model of regional adjustment to external trade liberalization. Section 3 describes the quasi-experimental empirical setting and the data. Our estimation strategy is described in Section 4, and empirical results are reported in Section 5. In Section 6, we examine the behavior of the theoretical model with a view to reproducing our key estimated parameter. Section 7 concludes.

2. Theory

2.1. A three-region geography model

Our theoretical starting point is the variant of Krugman's (1991) "new economic geography" model proposed by Helpman (1998), which offers an attractive framework for the analysis of market-access effects at the region level, as it explicitly considers congestion costs due to a non-tradeable resource H, thought of as housing.[2]

Details of the model are given in the Appendix. Here, we focus on sketching its main elements. The model features three regions, indexed by i: two regions in A(ustria) and one region R(est of the world). A is composed of an interior region I and a border region B. Labor, L, the sole production factor, is assumed to be fully employed and perfectly mobile within A but immobile between A and R. Workers spend a fraction μ of their income on varieties of a differentiated traded good, M, with a taste for variety represented by the substitution elasticity σ. The remaining fraction of income, $1-\mu$, is spent on housing H. The market for M is Dixit–Stiglitz monopolistically competitive. Individuals decide where to locate according to the indirect utility they obtain from consumption of M and H.

Our comparative-static exercise will consist of tracking changes in nominal wages and employment within A as trade costs between A and R are lowered. We are interested in the parameter ρ, the border region's differential change in employment, (β), relative to its differential change in the nominal wage, (α), induced by the fall in external trade costs. This elasticity represents the slope of the regional labor supply curve. A high value of ρ means that employment reacts strongly while nominal wages do not, implying a relatively elastic interregional labor supply; and vice-versa for a low value of ρ. As our simulations will show, ρ is not only a highly policy relevant variable but it also turns out to be robust to assumptions on trade costs and country sizes for which it is impossible to determine the "realistic" values.

The non-linearity of the model makes it algebraically unsolvable. We therefore resort to numerical simulations.[3]

2.2. The experiment

As we seek to model external trade liberalization of an integrated country, we assume low trade costs within A, and we let trade costs between B and R decline from an almost prohibitive level to the same low level that we assume to exist within A.

[1] For a survey of the literature on within-country spatial effects of international trade, see Brülhart (2011).

[2] Using this model will allow us to compare our results to those obtained by Redding and Sturm (2008). In Section 6, we shall extend the Helpman model by introducing heterogeneous locational preferences as an additional dispersion force.

[3] The Maple files used for the simulations are available from the authors. The model can in principle imply multiple and unstable equilibria. We have ascertained that the equilibria obtained for each set of parameter values are unique and stable. The uniqueness and stability condition for equilibria in the Helpman (1998) model is $\sigma(1-\mu)>1$. Some parameter combinations used in our simulations violate this condition. Nonetheless, the equilibria we obtain turn out to be stable and unique. The reason is that, in our three-region version of the Helpman model, only a fraction of world demand is mobile (regional demand within A). Therefore, forces that favor instability are attenuated compared to the original two-region model. The extended version of this model (Section 6) is more stable still than the baseline model, since it contains an additional dispersion force in the form of taste heterogeneity.

Regions are separated by iceberg trade costs, such that for every unit sent from region i to region j only a fraction $\tau_{ij} \in (0,1)$ arrives in j. The geographical structure of the three-region model is represented by the following assumptions on trade costs:

$$\tau_{IR} = \tau_{IB}\tau_{BR},$$

which means that for a unit of the M-good to be transported between I and R it has to transit through B. Thus, the border region is nearer to R than the interior region.

We choose the following parameter values to simulate external trade liberalization:

$$\tau_{IB} = 0.9,$$
$$\tau_{BR} = \{0.1, 0.2, 0.3, 0.4, 0.5, 0.6, 0.7, 0.8, 0.9\}.$$

We solve the model for each of the nine levels of τ_{BR}, and we compute the percentage change in equilibrium nominal wages, $\dot{w}_i$, and employment, $\dot{L}_i$, for each 0.1 increment of trade cost reduction.[4]

We can then calculate the ratio between the difference in growth rates of employment and the difference in growth rates of wages:

$$\rho \equiv \frac{\dot{L}_B - \dot{L}_I}{\dot{w}_B - \dot{w}_I}.$$

This ratio is computed for every increment of trade-cost reduction, which yields eight such ratios for each combination of parameters other than τ_{ij}. As we will show, it turns out that ρ varies only trivially across pairs of trade costs for which it is calculated. We will therefore report averages of the eight computed ratios.

To calibrate this model, we need to decide on the values of the following parameters: housing stocks (in each region), H_i, population in A and R, the elasticity of substitution among differentiated goods, σ, and the expenditure share of housing, $1-\mu$. The population distribution within A is, of course, endogenous.

In order to cover the range of recent empirical estimates of substitution elasticities, we experiment with values of σ in the interval from 3 to 6.[5] As we shall see, the value assumed for the housing share $(1-\mu)$ is crucial. We take 0.25 as our best guess but shall explore the implications of alternative values. According to the OECD input–output table for Austria in 1995, housing expenditure amounted to 25% of the total wage bill and of 15% of the total wage bill plus net profits.[6] The distribution of housing stocks within A is obtained by calibrating the model so as to replicate the population distribution observed in our data.[7] We exogenously assign a distribution of the total stock of housing between A and R, choosing $H_R = \frac{H}{3}$ and $L_R = \frac{L}{3}$ and normalizing total stock of housing and labor by setting $H = L = 1$. Hence, A is twice the size of R. This is arbitrary, but, as we shall show in Section 6.1, the implied ρ s are almost unaffected by different parametrizations of H_R and L_R as well as to different-sized changes in trade costs.

[4] $\dot{w}_i$ and $\dot{L}_i$ are percentage changes between steady states. To be clear, let $w_{\tau_{BR}=0.1}$ and $w_{\tau_{BR}=0.2}$ be equilibrium wages in B when $\tau_{BR}=0.1$ and when $\tau_{BR}=0.2$, respectively. Then $\dot{w}_B = (w_{\tau_{BR}=0.2} - w_{\tau_{BR}=0.1})/w_{\tau_{BR}=0.1}$, and analogously for wages in I and for employment. The empirical counterparts are cumulative growth rates over the entire pre- and post-liberalization subperiods, assuming that these subperiods are sufficiently long to capture the full transition between steady states.

[5] See, e.g., Baier and Bergstrand (2001), Bernard et al. (2003), Hanson (2005), Broda and Weinstein (2006) and Head and Mayer (2006).

[6] Davis and Ortalo-Magné (2011) find that, between 1980 and 2000, the median US household expenditure share of housing was a stable 0.24, with a standard deviation of 0.02.

[7] In our data set, municipalities belonging to our baseline definition of the border region (B) accounted for 5.1% of Austrian population prior to liberalization. Their implied housing stock in our calibrations ranges from 6 to 9% of the total for country A.

Table 1
Simulated ρ — baseline model.

	$\sigma=3$	$\sigma=4$	$\sigma=5$	$\sigma=6$
$(1-\mu)=0.20$	10.33	9.60	9.23	9.00
$(1-\mu)=0.25$	7.70	7.16	6.88	6.71
$(1-\mu)=0.30$	5.97	5.54	5.33	5.20
$(1-\mu)=0.40$	3.82	3.55	3.43	3.33
$(1-\mu)=0.50$	2.54	2.36	2.27	2.21

Note: Reported numbers are simulated equilibrium values of ρ, the measure of employment adjustment relative to wage adjustment.

2.3. Simulation results

Table 1 reports the simulated values of ρ for several combinations of σ and $(1-\mu)$. The values of this ratio range from 2.21 to 10.33. For what we consider our most realistic parameter combination, $\sigma=4$ and $(1-\mu)=0.25$, the predicted ρ equals 7.16. This implies that the magnitude of trade-shock induced employment growth in the border region is some seven times larger than the magnitude of the trade-shock induced increase in nominal wages. At face value, this could be taken to suggest rather elastic interregional labor supply.

As a check on the robustness of this result, we computed implied values of α, β and ρ for different levels and changes of external trade costs and for different relative sizes of country A and the rest of the world B.[8] The implied wage and employment effects turn out to be sensitive to these assumptions: the larger the cut in external trade costs, and the larger the size of the outside economy, the larger are the simulated values of α and β. This is why looking at these effects themselves would be of little help in mapping the model to the data. When we focus on their ratio, however, this issue no longer arises, as ρ turns out to be robust to modeling choices on variables other than σ and $(1-\mu)$.[9] This lack of sensitivity is not surprising. By increasing the size of R, for instance, trade liberalization becomes more important for both I and B, but more so for B. Yet, ρ is not a measure of the locational attractiveness of B relative to I; rather, it captures whether that increased attractiveness manifests itself more in terms of employment growth or in terms of nominal wage growth. This ratio is largely insensitive to the overall attractiveness of B with respect to I.

We now turn to an empirical estimation of ρ.

3. Empirical setting and data

3.1. Austria and Eastern Europe before and after the fall of the Iron Curtain

The experience of Austria over the last three decades provides a propitious setting, akin to a natural experiment, within which to explore regional responses to changes in trade openness. In 1975, at the beginning of the period covered by our study, Austria lay on the eastern edge of democratic, market-oriented Europe. By 2002, which marks the end of our sample period, it found itself at the geographical heart of a continent-wide market economy. We argue that the fall of the Iron Curtain can be thought of as an exogenous change in market access, that it was unanticipated, that it was large, and that it affected different Austrian regions differently.

We assume that the lifting of the Iron Curtain was exogenous to events in Austria. Moreover, during the period covered by our study, this transformation took the form of a trade shock: a large change in

[8] The results of these simulations are shown in Table A1 of the online "supplementary materials".

[9] In addition to the sensitivity analyses reported in Table A1 of the "supplementary materials", we have explored the implications of changing the assumed intra-country trade cost τ_{IB}. We found the simulated values of ρ to be essentially insensitive to this assumption as well.

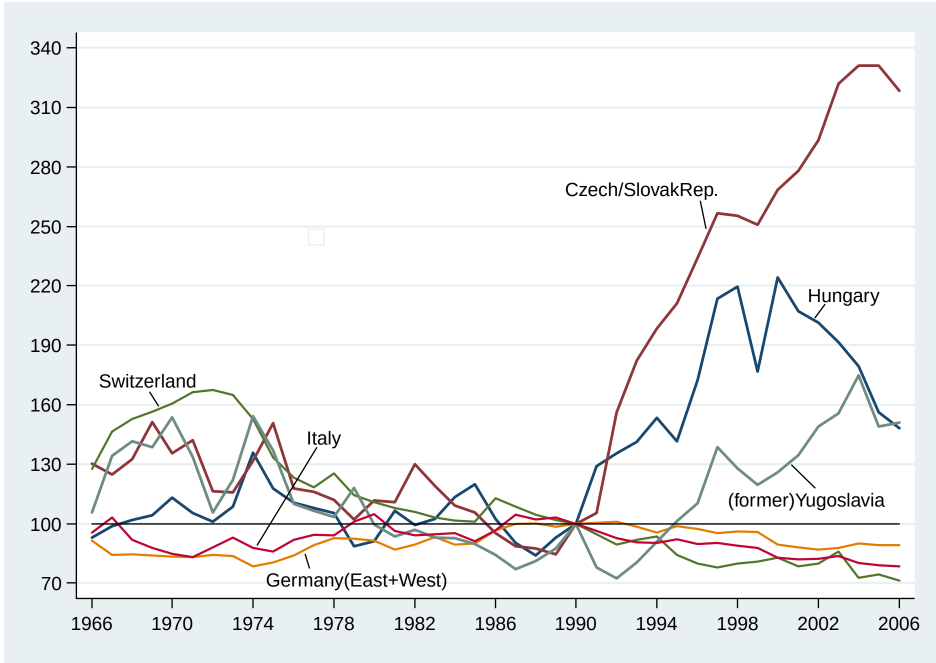

Fig. 1. Austria's Post-1990 eastward trade opening. Value of merchandise imports + exports as share of total Austrian trade, 1990 = 100. Source: UN Comtrade database.

cross-border openness of goods markets with little concomitant change in openness to cross-border worker flows.[10]

The timing of the main "exogenous shock" is also straightforward to pin down. While some economic reforms had started across communist Europe soon after the ascent of Mikhail Gorbachev in 1985, the rapid break-up of the Soviet bloc in 1989–90 took most contemporary observers by surprise. In January 1989, the fact that a mere two years later all of Austria's Comecon neighbors (Czechoslovakia, Hungary and Slovenia) as well as nearby Poland and even the Soviet Union itself would have held democratic elections and jettisoned most aspects of central planning, was unexpected by most.[11] Hence, we define 1990 as the watershed year 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, only fell gradually post-1990. The main milestones in this respect were the entries into force of free trade areas between the EU and Hungary, the Czech Republic, Slovakia and Poland in 1992, and with Slovenia in 1996.[12] Furthermore, the Eastern European countries all applied for full EU membership in the mid-1990s.[13] Austria itself had lodged its membership application in 1989 and joined the EU in 1995. In short, the decade following 1990 was a period of gradual but profound and lasting mutual opening of markets, to an extent that up to the very late 1980s had been largely unanticipated.

The magnitude and time profile of the post-1990 transformation can be gleaned from Fig. 1, where we report Austrian bilateral trade volumes with its neighboring countries, scaled relative to their 1990 values. The take-off in 1990 of trade between Austria and its formerly communist neighboring countries is evident. While, over the 1990s, the share of Austria's trade accounted for by its western neighbor countries shrank by between 13% (Germany) and 20% (Switzerland), it increased by 107% with Hungary and by 178% with the Czech and Slovak republics. Fig. 1 shows that trade with the former constituent parts of Yugoslavia only took off by the middle of the decade, which is unsurprising given the wars in Croatia and Bosnia–Herzegovina that lasted until 1995. Trade with Slovenia has been recorded separately since 1992. It shows a continuous increase as a share of Austrian trade of 78% between 1992 and 2002. The data thus confirm that 1990 marked the start of a large and sustained eastward reorientation of Austrian trade.[14]

[10] Free East–West mobility of workers only started to be phased in after EU enlargement in 2004, well after the end of our sample period. In a review of pre-enlargement migration patterns and policies, the OECD (2001) concluded that "except for Germany, the employment of nationals of the CEECs in OECD member countries did not increase significantly [post-1990]" (p. 35) and that "the current state of integration between the CEECs and the EU is characterized by limited labor flows but strong trade integration and increasing capital market integration" (p. 107). Austria had experienced considerable inflows of mainly fixed-term "guest workers" from Yugoslavia already before 1990. Available data from the WIFO's "SOPEMI Reports" show that the number of Yugoslav and CEEC workers in Austria in fact shrank between 1992 and 2001, from 134,000 to 71,000 and from 42,000 to 38,000 respectively. The treatment we analyze can therefore be considered as a trade shock. For an analysis of a cross-border opening of labor markets, see Buettner and Rincke (2007), who used German reunification as a quasi-experiment to explore the impact of migration on border-region employment and wages.

[11] Some quotes from *The Economist* magazine illustrate this point. In its issue of 7 January 1989 (p. 27), *The Economist* wrote of Gorbachev's "chance to relaunch [his] reforms for the start of the next five-year plan in 1991" but warned that "real reform […] may have to wait until the 1996–2000 plan". The centrally planned economy was evidently expected to last at least for the rest of the decade. In its 11 March edition (p. 14), *The Economist* speculated about a possible loss of power by Gorbachev and concluded that "if there were a bust-up over reform, the regime that would replace Mr Gorbachev's would probably be conservative, disciplinarian and much less interested in rejoining the world". This shows that informed opinion in early 1989 considered a continuation of the gradual Gorbachev reforms as the most likely (or even only) path towards East–West integration — with a considerable risk of a restoration of hardline communist control and the attendant economic isolation. A sudden collapse of the communist system did not feature among the scenarios considered probable until the second half of 1989, in particular after the fall of the Berlin Wall on 9 November of that year.

[12] Formally, these are the starting dates of "Interim Agreements". The official "Europe Agreements" entered into force two to three years later. Trade barriers were phased out gradually over up to ten years, but liberalization already started during the Interim Agreement period.

[13] Hungary and Poland applied in 1994, Slovakia in 1995 and the Czech Republic and Slovenia in 1996.

[14] The geographic reorientation of Austrian trade occurred against a background of steadily increasing overall trade orientation. Imports and exports corresponded to 58% of Austria's GDP in 1975, to 73% in 1989 and to 93% in 2002 (OECD data). This was a faster expansion than the OECD average (1975 definition): Austrian trade accounted for 1.43% of OECD trade in 1975, for 1.59% in 1989 and for 1.80% in 2002.

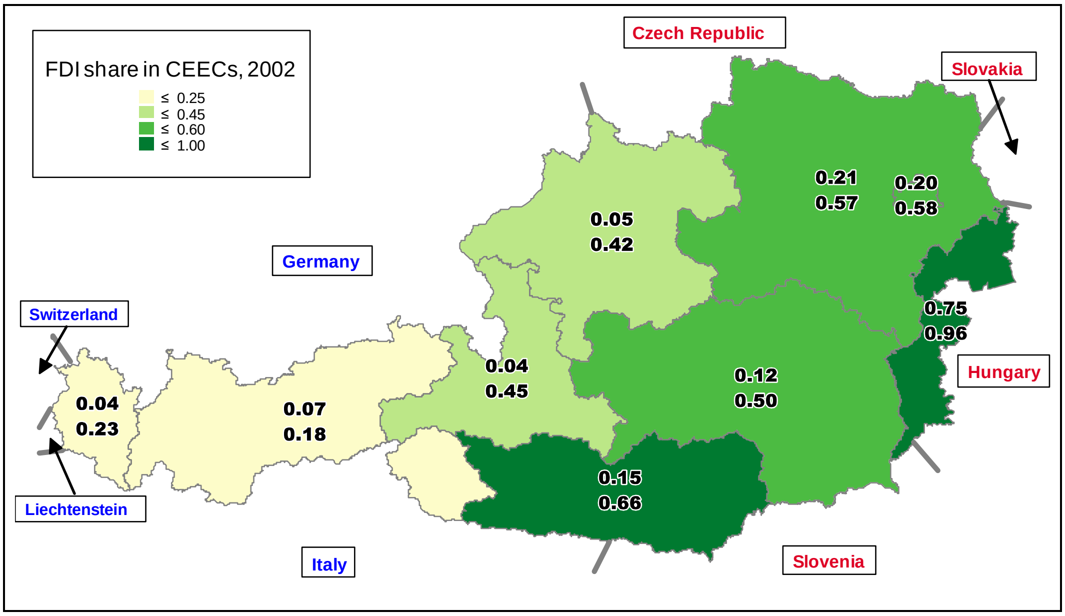

Fig. 2. Eastward orientation of Austrian regions. (Share of total outward FDI located in CEECs, top (bottom) numbers refer to 1989 (2002).) *Notes*: shares in terms of numbers of firms; CEECs comprise Albania, Belarus, Bosnia–Herzegovina, Bulgaria, Croatia, Czech Republic, Hungary, Estonia, Latvia, Lithuania, Macedonia, Moldova, Poland, Romania, Russia, Serbia–Montenegro, Slovakia, Slovenia, Ukraine.
Source: Austrian National Bank.

Austria's small size implies that access to international markets is important: it was the OECD's fifth most trade oriented country in 1990.[15] Moreover, simple inspection of a map reveals that the transformations in Austria's eastern neighbors should have affected Austrian regions with different intensity (see Fig. 2). Austria's East–West elongated shape 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. Regional trade data would allow us to corroborate this claim explicitly. No such statistics exist for Austria, but there is strong evidence from other countries of gravity-type trade patterns also at the sub-national level.[16] Furthermore, we can draw on region-level data on foreign direct investment (FDI) collected by the Austrian central bank. In Fig. 2, we report the CEEC share of the stock of outward FDI projects by Austrian firms. This map shows that firms in eastern Austria are significantly more oriented towards the eastern European markets than firms based in western Austria, and that this gradient has remained just as strong in 2002 as it was in 1989. The FDI data corroborate the trade data in showing how strongly the Austrian economy turned eastwards post-1990: the share of Austrian FDI projects hosted by CEECs rose from 14% of total Austrian FDI in 1989 to 51% in 2002. In 2002, a full 96% of FDI from Austria's most easterly region (Burgenland) was targeted at CEECs, while the corresponding share of Austria's most westerly region (Vorarlberg) was 23%. Austria thus provides us with considerable variation for identifying effects that are specifically due to improved access to eastern markets.

As we couch our analysis within a market-based model of spatial wage and employment adjustments, we need to ascertain that such a model is indeed appropriate for our empirical setting. Almost all Austrian firms are bound by industry-level collective wage agreements. These agreements allow for some regional differentiation. More important, however, is the fact that the agreed rates serve as wage floors that are rarely binding and thus allow for considerable flexibility across firms and regions. In 2001, for example, the average agreed wage rate in the highest-wage region (Vorarlberg) exceeded that of the lowest-wage region (Burgenland) by 17%, and the corresponding difference in average effective wage rates amounted to fully 36%.[17] Another piece of evidence of relatively flexible private-sector wage setting in Austria is given by Dickens et al. (2007), who show that in a sample of 16 industrialized countries, Austria has the seventh-lowest downward rigidity of nominal wages - somewhat more rigid than the UK, but somewhat less rigid than Germany and considerably less so than the United States. We conclude that Austria provides an appropriate setting for our analysis also in terms of the structure of its labor market.

3.2. A data set on wages and employment in Austrian municipalities

Our analysis is based region-level measures of employment and wages computed from the Austrian Social Security Database (ASSD). The ASSD records individual labor-market histories, including wages, for the universe of Austrian workers.[18] 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 midpoint of each quarter, yielding 112 measurements from the first quarter of 1975 to the fourth quarter of 2002.

The wage data are right censored, because social security contributions are capped at a level that is adjusted annually, and effective income exceeding that limit is not recorded. In order to minimize distortions from such censoring, we construct wages as medians across individuals by municipality.[19] Wages are recorded on a per-day basis,

[15] Only Luxembourg, Belgium, Ireland and the Netherlands had higher trade-to-GDP ratios.
[16] See, for example, Combes et al. (2005) and Helble (2007) for Europe, and Hillberry and Hummels (2003, 2008) for the United States.
[17] These data are taken from the 2002 statistical yearbook of the Austrian Federal Economic Chamber.
[18] For a thorough description, see Zweimüller *et al.* (2009). Public-sector workers are not covered by this database prior to 1988, nor are the self-employed. We therefore work exclusively with data pertaining to private-sector employees.
[19] A comparison of annual median wages (reported by Statistics Austria) to the censoring bounds in the ASSD (reported by Zweimüller *et al.*, 2009), shows that the former falls very comfortably between the latter in all our sample years.

which means that they are broadly comparable irrespective of whether employment contracts are part-time or full-time.

The ASSD assigns every establishment to one of 2305 municipalities. Our identification strategy will hinge on the relative distances of these municipalities to eastern markets. Our main measure is the road distance to the nearest border crossing to one of Austria's formerly communist neighbor countries. As an alternative, we use the shortest road travel time between each municipality and the nearest eastern border crossing, computed as road distances weighted by average traveling speeds.[20] Since we can allocate firms to one of 16 sectors, we can furthermore control for the industrial composition of municipalities.[21]

4. Estimation strategy

Our basic estimation strategy follows the difference-in-difference approach applied by Redding and Sturm (2008). We regress the endogenous variable of interest on the interaction between a dummy for border regions (*Border*) and a dummy that is equal to one for all years from 1990 onwards (*Fall*), as well as on a full set of time (t) and location (i) fixed effects. The coefficient estimated on the interaction term measures whether and how the dependent variable evolved differently in border regions (the treatment group) compared to interior regions (the control group) after the fall of the Iron Curtain.

Specifically, we estimate the following equation for median nominal wage growth:

$$\Delta Wage_{it} = \alpha(Border_i \times Fall_t) + d_i + d_t + \varepsilon_{it}^{wage}, \quad (1)$$

where, in our baseline specification, $\Delta Wage_{it}$ is the annual growth rate measured at quarterly intervals:

$$\Delta Wage_{it} = \frac{Wage_{it} - Wage_{it-4}}{[Wage_{it} + Wage_{it-4}]^{*}0.5},$$

d_i denotes a full set of municipality fixed effects, d_t denotes a full set of quarter fixed effects, and ε_{it}^{wage} is a stochastic term. Unobserved time-invariant heterogeneity in municipal wage levels is differenced out by taking growth rates. Furthermore, the municipality-specific dummies control for any unexplained differences in linear wage trends, and the time dummies control for nation-wide temporary shocks to median wage levels including the common impact of the fall of the Iron Curtain on median wages across all of Austria.[22]

We then apply a corresponding specification for changes in municipal employment:

$$\Delta Empl_{it} = \beta(Border_i \times Fall_t) + d_i + d_t + \varepsilon_{it}^{empl}, \quad (2)$$

where *ΔEmpl* is defined equivalently to *ΔWage*.

In an alternative specification, we express *ΔEmpl* and *ΔWage* as changes over the full pre- and post-1990 sample periods.

Our coefficients of interest are $\hat{\alpha}$ and $\hat{\beta}$. They capture the differential post-1990 trajectories of nominal wages and employment in border regions, which we interpret as the effect of increased market access subsequent to the fall of the Iron Curtain.

The ratio of the two coefficients, $\hat{\rho} = \frac{\hat{\beta}}{\hat{\alpha}}$, provides us with a measure of the relative magnitudes of employment and nominal wage adjustments, and thus of the slope of the average municipal labor supply curve, which we can compare to the value predicted by theory.[23]

As a complement to parametric estimation, we report non-parametric evidence on the relationship between, on the one hand, the growth of median wages or total employment in each municipality and, on the other hand, the distance of the respective municipalities to the eastern border. Specifically, we estimate the following equations:

$$\Delta Wage_{it} = \gamma_i(Fall_t \times d_i) + d_i + d_t + \omega_{it}^{wage}, \text{ and} \quad (3)$$

$$\Delta Empl_{it} = \delta_i(Fall_t \times d_i) + d_i + d_t + \omega_{it}^{empl}. \quad (4)$$

The parameters $\hat{\gamma}_i$ and $\hat{\delta}_i$ represent municipality-specific estimates of differential average growth after 1990 compared to the pre-1990 period. A plot of the relationship between these parameters and municipalities' distance to the eastern border can give us an indication of the market-access effect without any prior restriction on the definition of the treatment sample (i.e. of "border" municipalities).

Specifications (1) and (2) allow us to estimate treatment effects averaged over the full treatment period covered by the sample (1990–2002). One of our aims being to explore the time profiles of adjustment, we also estimate treatment effects separately for each year of the treatment period. We therefore also consider the following specifications:

$$\Delta Wage_{it} = \alpha_t\left(Border_i \times Fall_t \times d_t^{year}\right) + d_i + d_t + \nu_{it}^{wage}, \text{ and} \quad (5)$$

$$\Delta Empl_{it} = \beta_t\left(Border_i \times Fall_t \times d_t^{year}\right) + d_i + d_t + \nu_{it}^{empl}, \quad (6)$$

where d_t^{year} denotes year dummies. This gives us annual treatment effects $\hat{\alpha}_t$ and $\hat{\beta}_t$ for each year subsequent to the fall of the Iron Curtain.

Finally, we seek to control for the possibility that border regions differ systematically from interior regions not only in terms of geography but also in terms of size and industrial composition. We therefore reduce the set of control (interior) municipalities to those that provide the nearest match to at least one of the treatment (border) municipalities in terms of the sum of squared differences in sectoral employment levels, measured in 1989. We compute estimates of α and β as average treatment effects in a setup 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, since including municipality fixed effects may not account for all plausible covariance patterns (Bertrand et al., 2004). Hypothesis tests on $\hat{\rho}$ are Wald tests using the delta method to approximate the variance of $\hat{\rho}$, and taking account of the municipality-level clustering of the coefficient standard errors.

5. Results

5.1. Baseline empirical specification

For our baseline results, we define *Border* as comprising all municipalities whose geographic center is at most 25 road kilometers away from the nearest eastern border crossing, and "eastern" is defined as comprising all four formerly planned economies adjacent to Austria (Czech Republic, Hungary, Slovakia and Slovenia). A map of these municipalities is given in Fig. 3.

The raw data show that border municipalities had relatively low wages and were comparatively small in employment terms

[20] Road distances and travel times were obtained from Digital Data Services GmbH, Karlsruhe, Germany. 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.

[21] The list of sectors covers the full spectrum of economic activities and primarily consists of aggregates of NACE two-digit industries (see Zweimüller *et al.*, 2009).

[22] The main effects of $Border_i$ and $Fall_t$ are not identified due to the inclusion of d_i and d_t.

[23] Since our two estimating equations feature identical sets of regressors, estimating them separately by OLS is equivalent to estimating them as a system. Our strategy thus amounts to estimating the slope of the regional labor supply curve, $\hat{\rho}$, via indirect least squares.

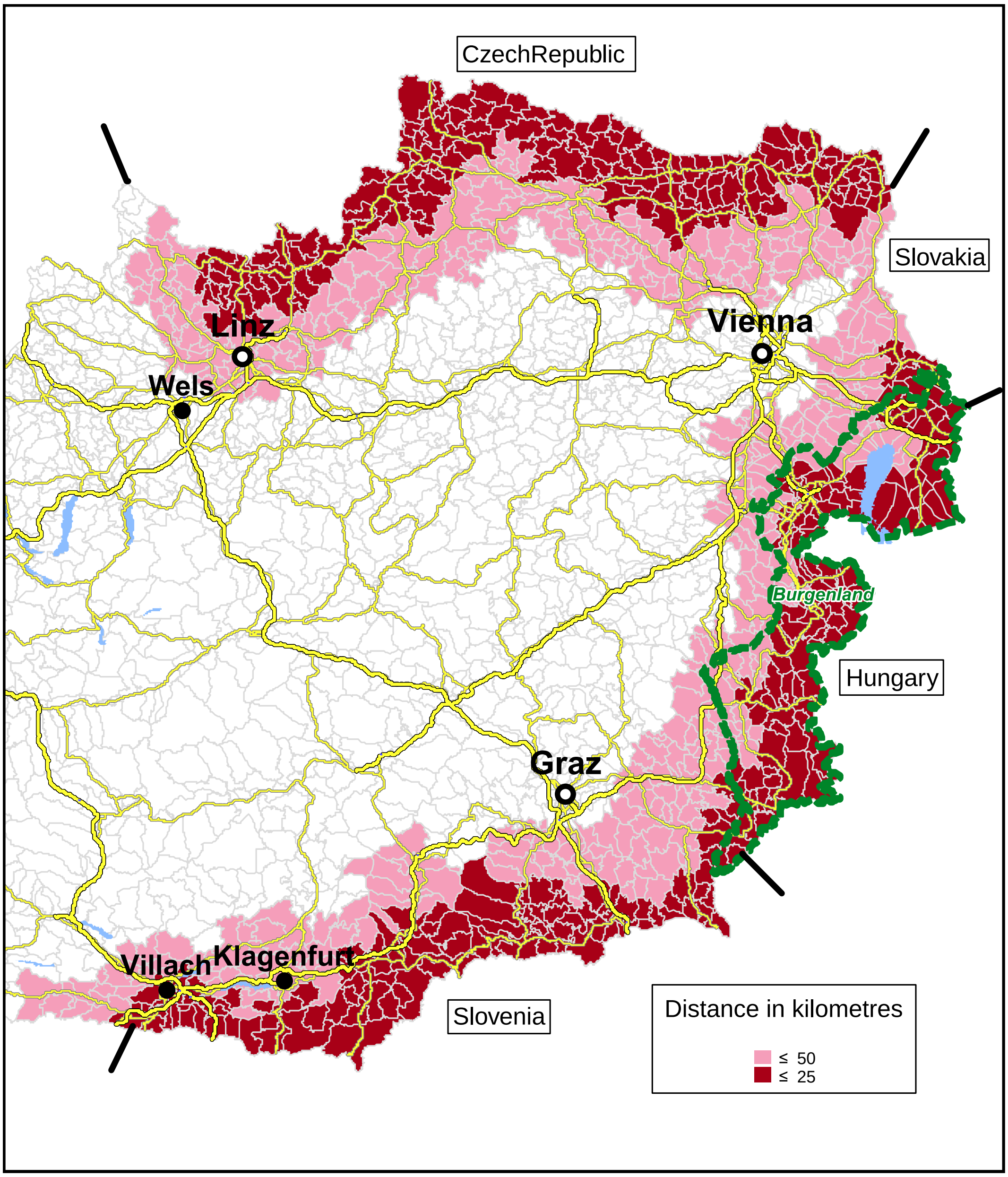

Fig. 3. Treatment municipalities.

throughout the period covered by the data.[24] Such differences in levels could be explained by a multitude of factors that it would be difficult to control for comprehensively. The same is true for changes over time across all municipalities: why some municipalities on average grow faster than others could be due to a range of variables it again would be impossible to capture in its entirety. This is why we focus on differences in changes pre- and post-1990 between border and interior regions. No major shock coincided with that timing and geographic reach other than the opening of the Eastern markets.[25]

Our baseline econometric estimates are shown in Table 2. Column 1 reports the coefficient $\hat{\alpha}$ from an estimation of the wage Eq. (1). The estimated coefficient implies that over the 13 years subsequent to the fall of the Iron Curtain, nominal wages grew 0.27 percentage points faster annually in border regions than in interior regions, relative to their respective pre-1990 growth rates. This effect is statistically significant at the five-percent level. It suggests that improved market access after the opening of Eastern markets has boosted nominal wages in the most affected Austrian municipalities. The corresponding estimate for employment growth, the coefficient $\hat{\beta}$ from an estimation of Eq. (2), is given in column 2 of Table 2. We again find a positive impact. The treatment effect of improved Eastern market access on the relative employment growth of border relative to interior regions is estimated as 0.86 percentage points, which is statistically significant at the one-percent level. In cumulative terms, our benchmark parameter estimates imply that, thanks to the opening of the Central and Eastern European markets, Austrian border regions experienced an approximately 5% increase in nominal wages, and a 13% increase in employment, relative to regions in the Austrian interior.[26]

[24] For summary statistics, see Table A2 in the online “supplementary materials”.

[25] One potentially confounding event was the eligibility of the Burgenland region for EU regional funds from 1995 onwards. We control for this in the robustness section, and find it to have no significant effect.

[26] The cumulative wage effect is calculated as $100\left(\left(1+\Delta w_{I,Fall}+\hat{\alpha}\right)^T-\left(1+\Delta w_{I,Fall}\right)^T\right)$, where $\Delta w_{I,Fall}$ is the median post-1990 growth rate of interior-region wages (= 3.56%, see Table A2 in the online “supplementary materials”), and T is the number of post-1990 sample years (= 13). The cumulative employment effect is calculated identically, *mutatis mutandis*.

Table 2
Baseline regressions. *Border* defined as municipalities within 25 km from Czech, Hungarian, Slovakian or Slovenian border.

Dependent variable:	Annual growth rate, quarter by quarter		Average annual growth rate, pre- and post-1990	
	ΔWage	ΔEmpl	ΔWage	ΔEmpl
	(1)	(2)	(3)	(4)
Border × Fall, 1990–2002	**0.267****	**0.861*****	**0.263*****	**0.803*****
	(0.12)	(0.28)	(0.07)	(0.31)
No. obs.	248,940	248,940	4610	4610
No. municipalities	2305	2305	2305	2305
R^2	0.058	0.021	0.049	0.021
Quarter fixed effects	Yes	Yes	–	–
Dummy for Fall, 1990–2002	–	–	Yes	Yes
Municipality fixed effects	Yes	Yes	Yes	Yes
Estimated rho	3.22		3.05	
	(1.79)		(1.45)	
H0: rho = 1 (*p* value)	0.213		0.157	
H0: rho = 3 (*p* value)	0.888		0.950	
H0: rho = 7 (*p* value)	0.035		0.007	

Note: estimation with OLS; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; *p* values on hypothesis tests based on delta method and clustered coefficient standard errors; *: p = 0.1, **: p = 0.05, ***: p = 0.01.

Our estimated coefficients $\hat{\alpha}$ and $\hat{\beta}$ suggest that trade liberalization has boosted wages as well as aggregate employment in Austrian border regions, but that the employment effect was some three times larger than the effect on wages (i.e. $\hat{\rho} = \frac{0.861}{0.267} = 3.22$). In this sense, employment was more responsive to changes in market access than nominal wages. The three tests shown in the bottom rows of Table 2 suggest that we can reject the hypothesis that $\hat{\rho} = 7$, as implied by the theoretical model of Section 2, but not that $\hat{\rho} = 3$, nor in fact that $\hat{\rho} = 1$.

Columns 3 and 4 of Table 2 show corresponding estimates with the respective dependent variables defined as average annual changes over the entire pre- and post-1990 sample periods. This reduces the sample size but changes the results only trivially. The implied value of $\hat{\rho}$ is very similar, with a point estimate of 3.05, and the hypothesis that $\hat{\rho} = 7$ is firmly rejected.

Our baseline point estimates of ρ are less than half as large as those implied by what we consider the most realistic calibration of the economic geography model of Section 2. If confirmed, this would represent a considerable divergence between theory and empirics. Before concluding that the model implies too much interregional labor mobility (i.e. too high a value of ρ), we therefore need to ascertain that our estimated value of $\hat{\rho}$ is a robust result.

5.2. Robustness

We begin by considering some alternative definitions of the treated region. In the first row of Table 3, we consider municipalities located between 25 and 50 km from the eastern border as a second treatment group. Our baseline estimates for the municipalities in the 0–25-kilometer range are robust to this additional control: they retain their magnitudes and statistical significance. Positive wage and employment effects are also found for the municipalities in the 25–50-kilometer range. However, the effects estimated for this outer band of border municipalities are only slightly more than half as large as those for the 25-kilometer border zone. Importantly, the estimated ratio $\hat{\rho}$, at 3.73, is close to the baseline estimate obtained for the 0–25-kilometer treatment group. Experimentation with even wider border definitions never yielded any statistically significant results. A corollary finding of our study, therefore, is that the regionally differentiated market access effects were confined to a rather narrow set of locations in close proximity of the border.[27]

In a second robustness test, we use an alternative distance measure: estimated road traveling time to the nearest official border crossing. This boils down to weighting roads by the speed at which they can be traveled. We report estimation results for a definition that attributes all municipalities located within 35 min from a border crossing to the treatment sample.[28] The results, shown in the second row of Table 3, are essentially equivalent to those of our baseline regressions.

As another manipulation of our basic setup, we drop Slovenia from the sample of relevant eastern markets. This has two reasons. One is that Yugoslavia, even though a centrally planned economy, was not a member of the Soviet bloc and was economically more open prior to 1990 than Austria's other eastern neighbor countries. The second reason is that the full potential of the Slovene market and those beyond it only emerged gradually over the 1990s, mainly as a result of the series of wars that accompanied the breakup of Yugoslavia.[29] We report these results in the third row of Table 3. When dropping Slovenia as a relevant eastern market, we find weaker evidence of a wage response and stronger evidence of an employment response among the municipalities in the reduced-size treatment group. However, these coefficients are very imprecisely measured, and we can reject none of the three hypotheses on $\hat{\rho}$.

In a second set of robustness checks, we consider alternative definitions of the control group. One potentially confounding feature of our empirical setting is the existence of Vienna - by far the largest Austrian city. Vienna is located 64 km, or 55 min, from the nearest eastern border (with Slovakia). It therefore is not included in our narrowly defined treatment groups. As it accounted for some 40% of Austrian employment in our data set overall, we nevertheless want to examine our baseline results against a specification that controls specifically for the 23 municipalities that constitute the city of Vienna. As can be seen in row 4 of Table 3, controlling for Vienna barely affects our baseline findings.

One might furthermore suspect some of our measured effects to be due to the region of Burgenland. As shown in Fig. 3, this region strongly overlaps with the set of municipalities defined as border regions with Hungary. Due to its relatively low per-capita income, Burgenland was granted Objective 1 status subsequent to Austria's accession to the European Union in 1995, making it eligible for generous regional subsidies. We therefore add a dummy variable that is equal to one for all observations that belong to Burgenland from 1995 onwards. These estimations are shown in the fifth row of Table 3. The inclusion of this control variable also has no significant effect on our coefficient estimates of interest.[30]

We next estimate our baseline models in samples of municipalities that are matched on industry-level employment. Thereby, we can examine whether our results might be driven by the fact that border municipalities happened to be specialized in sectors that experienced particularly pronounced growth after 1990. Rows 6 and 7 of Table 3 show average treatment effects of a matching estimator applied to differences in growth rates between the post-1990 and the pre-1990 periods. We match municipalities on employment levels in 16 industries. In row 7 of Table 3, we furthermore restrict the matched control municipalities to lie no closer than 70 km from the treatment municipalities. Since we match by the size of industries in terms of employment (and not in terms of employment shares), our matching strategy also controls for differences in the size of

[27] We provide further evidence of the steep spatial decay of the observed effects in Section 5.3.

[28] The overlap between the *Border* sample under the 25-kilometer definition and under the 35-minute definition is large but not perfect. The 35-minute sample encompasses 276 municipalities, of which 248 also feature in the 25-kilometer sample.

[29] Fig. 1 shows that Austrian trade with former Yugoslavia only took off around 1995 and did not expand to quite the same relative extent as trade with the three other Eastern neighbor countries.

[30] The coefficients on the Burgenland controls themselves, which we do not show in Table 3, are never statistically significant. Hence, Objective 1 status appears to have had no discernible impact on aggregate employment and wage growth in Burgenland.

Table 3
Robustness.

	Border × Fall, 1990–2002	Dependent variable:		Estimated rho	p values		
		ΔWage (1)	ΔEmpl (2)	(3) = (2)/(1)	Rho = 1 (4)	Rho = 3 (5)	Rho = 7 (6)
(1)	Border 0–25 × Fall, 1990–2002	**0.231*** (0.12)	**0.994***** (0.29)	**4.3** (2.59)	0.202	0.614	0.297
	Border 25–50 × Fall, 1990–2002	**0.143** (0.11)	**0.533**** (0.27)	**3.73** (3.31)	0.409	0.825	0.323
(2)	Border defined in terms of travel time	**0.271**** (0.13)	**0.877**** (0.41)	**3.24** (2.17)	0.303	0.913	0.084
(3)	Border defined without Slovenia	**0.150** (0.14)	**0.973***** (0.33)	**6.49** (6.28)	0.383	0.579	0.933
(4)	Controlling for Vienna × Fall	**0.269**** (0.12)	**0.913***** (0.28)	**3.39** (1.83)	0.191	0.830	0.049
(5)	Controlling for Vienna × Fall, and for Burgenland × post-1995	**0.284**** (0.13)	**0.884***** (0.30)	**3.11** (1.76)	0.229	0.950	0.027
(6)	Controlling for industrial composition: ATE, matching on 1989 employment in 16 sectors	**0.241**** (0.08)	**0.654**** (0.39)	**2.72** (1.87)	0.359	0.879	0.022
(7)	Controlling for industrial composition: ATE, matching on 1989 employment in 16 sectors and geographic constraint (70 km)	**0.247**** (0.09)	**0.934**** (0.41)	**3.78** (1.93)	0.150	0.686	0.095
(8)	Weighted Least Squares, baseline specification	**0.458***** (0.12)	**0.815***** (0.31)	**1.78** (0.83)	0.346	0.140	0.000
	No. obs.	248,940					
	No. municipalities	2305					
	quarter fixed effects	Yes					
	municipality fixed effects	Yes					

Note: estimation with OLS; standard errors in parentheses: heteroskedasticity consistent and adjusted for municipality-level clustering; *p* values on hypothesis tests based on delta method and clustered coefficient standard errors; *: p = 0.1, **: p = 0.05, ***: p = 0.01.

municipalities. Again we find statistically significant treatment effects on employment as well as on wages, and ratios of close to 3.

As a final check on our baseline results, we estimate specifications (1) and (2) using weighted least squares regression, taking sample-average municipal employment as weights, so as to reduce the weight of very small municipalities. As shown in row 8 of Table 3, our qualitative findings remain unchanged, but the magnitudes and statistical significance of the relevant coefficients increase. The wage effect is now statistically significant at the one-percent level as well, with the employment effect estimated to be only 1.78 times as large as the wage effect. Our baseline estimated values of the wage and employment effect, however, remain within the 95-percent confidence intervals also of these estimates.

For the eight specifications reported as robustness tests, we obtain estimated ratios of employment to wage adjustment, $\hat{\rho}$, ranging from 1.78 to 6.49 (Table 3, column 3). The hypothesis tests shown in columns 4 to 6 of Table 3 allow us to reject the hypothesis $\hat{\rho}=7$, which is implied by what we consider the most plausible calibration of the theoretical model of Section 2, in six of our eight runs. The hypothesis $\hat{\rho}=3$, however, is never rejected. Hence, the data do appear to point to relatively less quantity adjustment than predicted by the theory.

One aspect that our data do not allow us to control for is individual worker characteristics. We therefore cannot distinguish wage increases that are due to skill upgrading from wage increases that are due to higher wage premia for identically skilled workers. Recent work by Frías et al. (2009) suggests that differential industry-level trade-induced wage changes are explained almost entirely by wage premia, with no significant explanatory power for skill upgrading. Their result is based on Mexican data, where skill upgrading would appear a more likely adjustment channel than in Austria. Based on this evidence, skill upgrading does not appear as a likely unobserved confound biasing our results.

5.3. Non-parametric illustrations: space and time

So far, we have imposed a dichotomy between treatment (*Border* = 1) and control (*Border* = 0) municipalities. We now relax this by estimating specifications (3) and (4) and plotting the estimated post-1990 growth differential of each municipality against that municipality's distance from the eastern border.[31] Natural spline regressions shown in Figs. 4 and 5 respectively.[32] The plots show that there is a statistically significantly positive effect on both wages and employment for municipalities that are located close to Austria's eastern border, whereas there is none for municipalities beyond about 50 km from the border, with Vienna representing an evident outlier.

This representation confirms that the differential effect of post-1990 market opening was confined to a relatively narrow band of Austrian municipalities located close to the border. Our analysis corroborates the relatively sharp distance decay of intra-national market-access and agglomeration effects found elsewhere (see, e.g., Rosenthal and Strange, 2003).

Although the theory does not feature explicit dynamics, we consider it interesting to investigate the time profile of our estimated treatment effects. We can describe the disaggregate time profile within that period by estimating specifications (5) and (6). These regressions provide us with annual estimates of differential wage changes ($\hat{\alpha}_t$) and employment changes ($\hat{\beta}_t$) in border regions for each year post-1990. The results are shown in Figs. 6 and 7. In most sample years, border-region wage and employment growth rates did not diverge statistically significantly from those in interior regions. We do, however, observe two periods over which significant treatment effects are in evidence: in 1995–1997, border-region nominal wages exhibit significantly positive differential growth, and in 1997–2000 a corresponding spike is observed for border-region employment growth. Our results thus suggest that wages adjusted earlier than employment, which is consistent with the view that wages are quicker to react to changed market conditions (at least in upward direction) than employment levels. Note, however, that both responses occur with a lag of some five years after the fall of the Iron Curtain. This is likely due not only to sluggish market responses

[31] The raw scatter plot for wages is given in Figure A1 and that for employment is given in Figure A2 of the online "supplementary materials".

[32] The smoothed lines are obtained by creating variables containing a cubic spline with seven nodes of the variable on the horizontal axis (distance to the eastern border), and by plotting the fitted values obtained from an employment-weighted regression of the dependent variable (post-1990 growth wage/employment growth) on the spline variables.

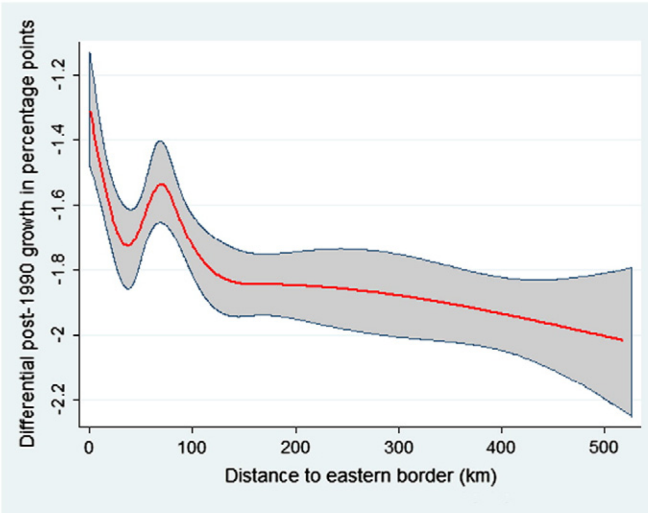

Fig. 4. Distance to border and Post-1990 wage growth – nonparametric fit. Natural spline regression on municipality-level differential post-1990 growth in median nominal wages ($\hat{\gamma}_i$ of Eq. 3).

but also to gradualism in the reduction of trade barriers and to persistence of political risk (with fears of a political backlash in Eastern Europe persisting well into the 1990s).

6. Revisiting the model

6.1. Allowing for preference heterogeneity

We find the magnitude of employment adjustment to equal around three times that of wage adjustment in our data — considerably lower than the ratio predicted by the most plausible calibration of the theoretical model of Section 2. Table 1 shows that, for the model to predict a ratio ρ of 3, we would need a housing share $(1-\mu)$ of between 0.4 and 0.5. This is too high to be realistic. We therefore conclude that the Helpman (1998) variant of the three-region economic geography model predicts too much employment adjustment and too little wage adjustment. For a better match between the theory and our empirical result, a stronger dispersion force is needed than that represented by housing alone.

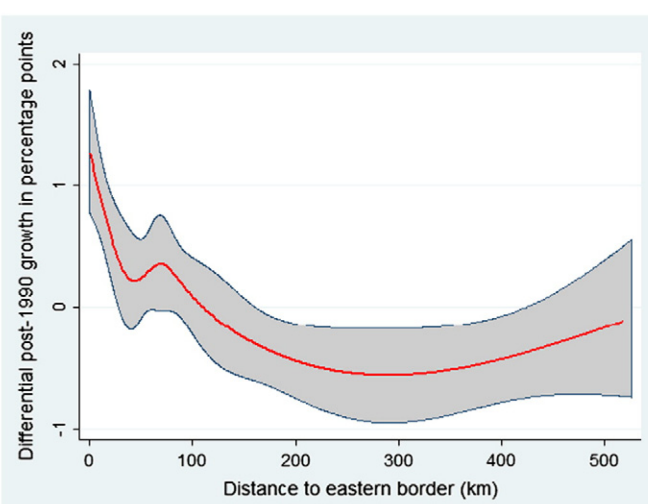

Fig. 5. Distance to border and post-1990 employment growth – nonparametric fit. Natural spline regression on municipality-level differential post-1990 growth in total employment ($\hat{\delta}_i$ of Eq. 4).

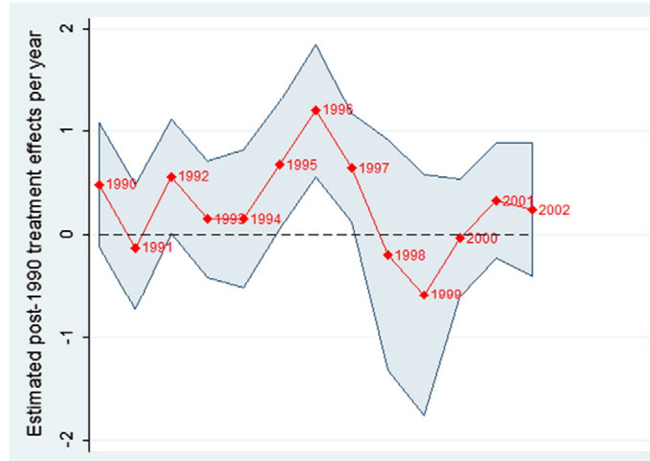

Fig. 6. Time profile of treatment effects — wages. Points correspond to $\hat{\alpha}_t$ of Eq. (5); bands represent 90% confidence interval.

We therefore consider a simple extension to the model by allowing for a plausible (though not the only conceivable) additional dispersion force: randomly distributed idiosyncratic locational preferences, following Tabuchi and Thisse (2002) and Murata (2003). Details of the model are again given in the Appendix. Preference heterogeneity is modeled through the parameter $\chi \in (0,\infty)$. When $\chi = 0$, individuals have identical preferences and choose their region of residence solely according to the indirect utility derived from their consumption of *M* and *H*. This is the preference structure of the model we considered in Section 2. As χ increases, idiosyncratic locational preferences become more important, and in the extreme case of $\chi \rightarrow \infty$ they alone determine workers' location choices.

There is neither empirical nor theoretical guidance as to what value to assign to χ. We will, however, be able to gauge the plausibility of values of χ indirectly. The presence of heterogeneity gives rise to regional real-wage differences that are not eliminated by migration precisely because, with heterogeneity, there will be some workers who prefer not to migrate despite thereby foregoing an increase in the real wage. We can thus assess values of χ by looking at the implied share of workers that do not move despite a given regional difference in real wages. For a plausibility check, we can draw on some related empirical evidence, based on the mobility of unemployed workers (see Shields and Shields, 1989, for an early survey). Faini et al. (1997) found that the percentage of Italian unemployed refusing

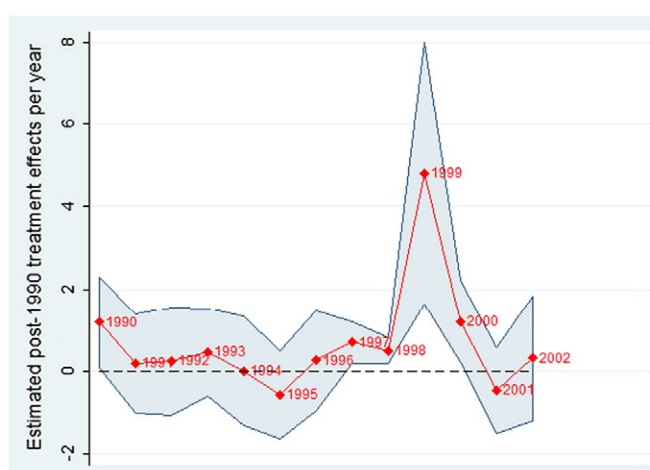

Fig. 7. Time profile of treatment effects — employment. Points correspond to $\hat{\beta}_t$ of Eq. (6); bands represent 90% confidence interval.

Table 4
Extended model: implied immobility for $\rho\approx3$.

	$\sigma=3$	$\sigma=4$	$\sigma=5$	$\sigma=6$
$(1-\mu)=0.20$	25	29	30	31
	[19]	[22]	[23]	[25]
$(1-\mu)=0.25$	24	28	31	32
	[14]	[16]	[17]	[18]
$(1-\mu)=0.30$	20	25	27	29
	[10]	[12]	[12]	[12]

Note: Reported numbers are implied percentage real-wage differentials between regions within country *A*, such that $\rho\approx3$. Numbers in brackets are implied shares of country *A*'s population that prefer not to migrate at the prevailing real-wage differential.

to move out of their town of residence if a job were available elsewhere ranges from 21% (Northern male university graduates) to 61% (Southern low-education females). Fidrmuc (2005) reported survey evidence according to which 34% of EU15 unemployed and 25% of Czech unemployed stated in 2002 that they would not move under any circumstances even if a job became available elsewhere. These studies point towards considerable locational inertia even within countries, supporting the relevance of incorporating factors other than wage differentials among the determinants of labor mobility in models of economic geography.

We allow χ to take any non-negative value, and search for the value of χ that yields an equilibrium ρ of 3.[33] For each of these simulations we report the implied interregional real-wage difference and the implied population share of non-movers at that real-wage difference. The combination of these numbers allows us to gauge the plausibility of the implied value of χ.[34]

The corresponding results are reported in Table 4. Each cell of that table shows the implied percentage real-wage differential between regions within country *A* and, in brackets, the implied share of country *A*'s population that prefers not to migrate at the prevailing real-wage differential. Table 4 shows that allowing for heterogenous locational preferences allows us to align the model's predictions with our estimated ρ. We consider eight parameter combinations for σ and $(1-\mu)$, taking what we deem the most plausible values of these parameters. In all eight cases, a relatively small amount of preference heterogeneity suffices to produce a predicted value of $\rho=3$. The necessary degree of preference heterogeneity when $\sigma=4$ and $(1-\mu)=0.25$, for instance, is such that 16% of the population would not move even if the real wage were 28% higher in the other region. In light of the available European evidence on the issue, this does not appear to be an excessive dose of assumed intrinsic insensitivity to regional wage differentials.

6.2. Discussion

Our simulations suggest that the baseline economic geography model with housing as the sole dispersion force implies more labor mobility than our empirical estimates, and therefore overpredicts the importance of the employment adjustment channel relative to the wage adjustment channel. If we extend the baseline model by including a moderate amount of locational taste heterogeneity, we can easily reconcile the theoretical model with the empirical estimates.

On the face of it, our central result therefore stands in contrast to the findings of Hanson (2005) and Redding and Sturm (2008), who both concluded that the calibrated Helpman (1998) model fit their empirical estimates well.

For parameter values in the same range as those used in our paper, Redding and Sturm (2008) found that the Helpman model can replicate the growth differential of small and large cities subsequent to the loss of access to eastern markets following the division of Germany. Their analysis concentrated on adjustment via factor quantities, measured by population, as wage data are not available for the long time period covered by their study. Our results suggest that their conclusions might have been different had they been able to consider wage data. To see this, consider for instance the ten combinations of σ and $(1-\mu)$ that Redding and Sturm (2008, Table 3) have identified as offering the best match between the model and their empirical estimates. In each case, we can apply these parameters to the unamended (Helpman) variant of the three-region model and indeed find levels of trade integration, τ_{BR}, for which the model precisely matches the estimated coefficient of the baseline employment regression, $\hat{\beta}=0.86$ (see Table 2). The implied values of ρ across these ten calibrations range from 3.2 to 11.8. Only two calibrations yield ρs below 4, and they both imply rather large housing shares (of 42 and 48% respectively). The parameter configurations in the plausible range, i.e. with housing shares below 0.3, all yield ρs in excess of 6. Hence, information on wage effects does appear to be important for a full evaluation of the congruence between the theory and the data.

The analysis by Hanson (2005) concentrated on adjustment via factor prices, by estimating a structural wage equation of the Helpman model on US county data. His estimations imply plausible parameter values, with predicted housing shares if anything on the low side.[35] A comparison of his results to ours thus suggests that obstacles to labor mobility, even at a small spatial scale, are higher in Europe than in the United States. The logical upshot is that, while a geography model with immobile housing and homogeneous locational tastes offers a good fit with observed spatial adjustment the North American context, an additional dispersion force, such as heterogeneous tastes, ought to be considered in a European setting.

This result has implications for policy. It is an additional piece of evidence pointing to relatively lower labor mobility in Europe than in North America, even within countries. Hence, trade and other shocks with regionally asymmetric effects can bring about greater intra-national spatial wage inequality in Europe than in North America. However, if trade liberalization benefits previously low-wage regions as in the case of eastern Austria, then it can act to reduce spatial inequality.

7. Conclusions

We have used the opening of Central and Eastern European markets after the fall of the Iron Curtain as a natural experiment of the effects of trade liberalization on regional wages and employment. Identification is achieved by comparing differential pre- and post-liberalization growth rates of wages and employment between, on the one hand, Austrian regions located close to the border to the formerly closed and centrally-planned eastern economies and, on the other hand, Austrian regions further away from the border.

We find that trade liberalization has had statistically significant differential effects on both nominal wages and employment of a rather narrow band of border regions. Most of the observed impact was confined to locations within 25 km of the border, and no statistically significant effects are found beyond a distance of 50 km.

The estimated effect on employment exceeds the estimated effect on nominal wages by a factor of around three. Over the entire post-Iron Curtain period, locations within 25 km of the border are estimated to have experienced a 5% increase in nominal wages and a 13% increase in employment, relative to regions in the Austrian interior.

[33] We stop the search loop at the first iteration that implies a value of ρ between 2.9 and 3.1.

[34] If, for instance, in order to obtain a ρ of 3, χ had to be such that the real-wage difference between regions were 200% and the immobile population share were 95%, then, given the low plausibility of such a configuration, we would conclude that taste heterogeneity is not a useful modeling feature for matching the theory to the facts. Conversely, to the extent that equilibrium real-wage differentials and immobile population shares look plausible, heterogeneity in locational tastes can be considered an empirically relevant addition to the model.

[35] Hanson's (2005) mean parameter estimates across the four reported variants of the instrumented regressions for 1980–90 are $(1-\mu)=0.21$ and $\sigma=2.12$ (Table 4).

Wages are found to have reacted earlier than employment, consistent with the view that wages rise more quickly than employment levels in response to increases in regional demand.

We then calibrated a standard economic geography model featuring immobile housing, and compared the implied predictions to our estimation results. This comparison suggests that the model somewhat overpredicts the relative magnitude of employment adjustment and thereby implies too much mobility. When augmented by heterogeneous locational preferences, which adds an impediment to employment adjustment, the model is easily able to replicate the estimated ratio of employment and wage adjustment.

Appendix A. Theoretical model

We use multi-region versions of a model that combines features of Krugman (1991), Helpman (1998), Tabuchi and Thisse (2002) and Murata (2003).[36]

A.1. Demand

The world economy consists of Λ regions and is populated by a given mass of individuals, L, indexed by k. We divide the set of all regions into two subsets, which we call "countries", A(ustria) and R(est of the world). For notational convenience, we assume that regions 1 to λ belong to country A, while the remaining regions belong to country R. Labor is mobile within countries but immobile between countries.

Each individual is endowed with one unit of labor, which is the only factor of production. Individuals derive utility from the consumption of goods as well as, potentially, from an exogenous and idiosyncratic preference parameter associated with individual regions.

The component of utility that is associated with consumption is modeled as a Cobb–Douglas combination of a CES (Dixit–Stiglitz) aggregate of varieties of a tradeable good, M, and consumption of a non-tradeable resource, H:

$$U = \left(C^M\right)^{\mu}\left(C^H\right)^{1-\mu}, \qquad 0<\mu<1.$$

Since H is a non-tradeable and exogenously given local resource, we refer to it as "housing", following Helpman (1998).

Trade among regions incurs costs of the conventional "iceberg" type, whereby for each unit of a variety sent from location i to location j only a fraction $\tau_{ij}\in(0,1)$ arrives at its destination. Trade within regions is free, $\tau_{ii}=1, \forall i$; and bilateral trade costs are symmetric, $\tau_{ij}=\tau_{ji}\ \forall i,j$. Utility maximization under the budget constraint gives individual demand functions, and aggregation over all residents of a region results in the following demand functions for any domestic and any imported variety of good M, respectively:

$$x_{ii}^d = (p_{ii})^{1-\sigma}\left(P_i^M\right)^{\sigma-1}\mu E_i,$$

$$x_{ji}^d = \left(p_{ji}\right)^{1-\sigma}\left(P_i^M\right)^{\sigma-1}\mu E_i, \tag{7}$$

where the first subscript refers to the region where the variety is produced and the second subscript refers to the region where the variety is consumed. Thus, x_{ii}^d denotes demand for locally produced goods, and x_{ji}^d denotes demand for imports from another region j. There is no need for a variety-specific subscript, since, as discussed below, all varieties in a given region will have the same equilibrium factory-gate price for sales of locally produced goods, p_{ii}, and for imports, p_{ij}. Total income equals total expenditure, E_i, of which a constant fraction μ is spent on the aggregate of M varieties. The price index for tradeables, P_i^M, takes the following CES form:

$$P_i^M = \left[\sum_{j=1}^{\Lambda} n_j\left(p_{ji}\right)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}, \tag{8}$$

where n_j denotes the number of varieties produced in region j, and Λ is the number of regions.

The stock of H in each region is constant. Therefore, given expenditure shares, the equilibrium price of H is given by:

$$P_i^H = \frac{(1-\mu)E_i}{H_i}. \tag{9}$$

Total expenditure is the sum of labor income and income from local housing services:

$$E_i = w_iL_i + P_i^H H_i = w_iL_i + (1-\mu)E_i = \frac{w_iL_i}{\mu}. \tag{10}$$

In our baseline model of Section 2 (as in Krugman, 1991; Helpman, 1998), the indirect utility of a region-i resident is given by the real wage in that region:

$$\omega_i \equiv \frac{w_i}{\left(P_i^M\right)^{\mu}\left(P_i^H\right)^{1-\mu}}, \forall i. \tag{11}$$

In our extended model of Section 6 (as in Tabuchi and Thisse, 2002; Murata, 2003), total indirect utility is given by the sum of indirect utility derived from consumption (common to all individuals in a given region) and utility derived from the idiosyncratic appreciation that each individual k associates with region i:

$$V_i^k = \omega_i + \xi_i^k.$$

ξ_i^k denotes a random variable that is identically and independently distributed across individuals according to a double exponential (Gumbel) distribution with zero mean and variance $\pi^2\chi^2/6$. Given this distribution, the probability that an individual will choose to reside in region i is given by the logit formula

$$Pr_i(\omega_i,\chi) = \frac{\exp\left(\frac{\omega_i}{\chi}\right)}{\sum_i \exp\left(\frac{\omega_i}{\chi}\right)}, \tag{12}$$

where the sum in the denominator is taken over all domestic locations (λ for country A, and $\Lambda-\lambda$ for country R). Expression (12) implies that $\lim_{\chi\to\infty} Pr_i(\omega_i,\chi) = \frac{1}{\lambda}$ (for country A), which means that when the distribution of idiosyncratic locational preferences has infinite variance each region within a country has the same probability of being chosen, independently of the indirect utility obtained from consumption. Conversely, $\lim_{\chi\to 0} Pr_i(\omega_i,\chi) = \frac{1}{\sum_{j=1}^{\lambda}\exp(\omega_j/\omega_i)}$, which means that, in the absence of preference heterogeneity, regions are chosen solely on the basis of the utility derived from consumption. Analogous expressions hold for regions in R, where $\lim_{\chi\to\infty} Pr_i(\omega_i,\chi) = \frac{1}{\Lambda-\lambda}$ and $\lim_{\chi\to 0} Pr_i(\omega_i,\chi) = \frac{1}{\sum_{j=\lambda+1}^{\Lambda}\exp(\omega_j/\omega_i)}$.

A.2. Supply

Production functions are assumed to be identical in every region and characterized by a fixed labor input $F>0$, and a constant variable

[36] Tabuchi and Thisse (2002) use linear demand functions while Murata (2003) uses Dixit–Stiglitz preferences.

input per unit of output a. Total labor input l required to produce x units of output is:

$$l = F + ax.$$

The product market is monopolistically competitive. Profit maximization, under the large group assumption, yields the following pricing rules for own-region and other-region sales:

$$p_{ii} = \left(\frac{\sigma}{\sigma-1}\right) a w_i, \tag{13}$$

$$p_{ij} = \left(\frac{\sigma}{\sigma-1}\right) \frac{1}{\tau_{ij}} a w_i. \tag{14}$$

Expressions (13) and (14) reflect the well-known result that monopolistic competition with Dixit–Stiglitz preferences implies identical markups across firms. The marginal cost of producing for another region (which includes transport cost) is $1/\tau$ times the marginal cost of producing for the local market. Therefore, $p_{ij}/p_{ii} = 1/\tau > 1$. Since production technology is identical across firms and all firms perceive the same elasticity of demand, the optimal price is identical across firms in the same region. Prices p_{ii} and p_{ij} will differ across regions if and only if wages differ across locations. Using the optimal prices in the free-entry (zero profit) condition yields the equilibrium output of each firm, which is identical across regions:

$$\bar{x} = \frac{F}{a}(\sigma-1). \tag{15}$$

A.3. Equilibrium in labor and goods markets

Equilibrium in the labor market requires that the local supply of labor, L_i, equals labor demand:

$$L_i = n_i(F + a\bar{x}) = n_i F\sigma. \tag{16}$$

Solving Eq. (16) for n_i shows that the number of varieties produced in each region is in fixed proportion to the population of that location:

$$n_i = \frac{L_i}{F\sigma}. \tag{17}$$

Product–market equilibrium requires equality of supply and demand for any variety of M produced in each region. The supply and demand functions for varieties of the same region turn out to be identical and, therefore, equilibrium in the market for any variety ensures market-clearing for all varieties produced in the same region. The equilibrium condition for any of the varieties in region i is:

$$p_{ii}\bar{x} = \sum_{j=1}^{\Lambda} \left(p_{ij}\right)^{1-\sigma} \left(P_j^M\right)^{\sigma-1} \mu E_j, \forall i. \tag{18}$$

By Walras' law, if there is equilibrium in $\Lambda-1$ markets (whichever they are), the remaining market is in equilibrium as well. The system of equilibrium conditions in goods markets is therefore composed of $\Lambda-1$ independent equations. Substituting the expressions for optimal prices (Eqs. (13)–(14)), the price index (Eq. (8)), total expenditure (Eq. (10)), the number of varieties (Eq. (17)), and equilibrium output of any variety (Eq. (15)) into Eq. (18), the system of market-clearing equations for a given variety of M becomes[37]:

$$1 = \sum_{j=1}^{\Lambda} \frac{\left(\tau_{ij}\right)^{\sigma-1} (w_i)^{-\sigma}}{\sum_{k=1}^{\Gamma} L_k (\tau_{ik})^{\sigma-1} (w_k)^{1-\sigma}} w_j L_j . i = 1, \ldots, \Lambda-1. \tag{19}$$

A.4. Spatial equilibrium

A spatial equilibrium is defined as a geographical distribution of the population $\{L_i\}$ such that the probability that a given region is chosen equals the number of individuals who actually have chosen that region (Miyao, 1978). This definition is equivalent to the condition that in equilibrium net migration flows be zero (Tabuchi and Thisse, 2002). Thus, a spatial equilibrium requires the following:

$$L_A Pr_i(\omega_i, \chi) = L_i, \qquad \text{for } i = 1, \ldots, \lambda-1, \tag{20}$$

$$L_R Pr_i(\omega_i, \chi) = L_i \qquad \text{for } i = \lambda+1, .., \Lambda. \tag{21}$$

Since probabilities and populations sum to one in both countries, there is one less independent equation per country than there are regions.

Replacing Eqs. (8), (9), (10), (13), (14) and (17) into expression (11) and then replacing the resulting expression for real wages in i into Eqs. (20) and (21), we can rewrite Eqs. (20) and (21) as follows:

$$\frac{\exp\left(\dfrac{w_i/\chi}{\left(\sum_{j=1}^{\Lambda} \frac{L_j}{\sigma F} \left(\tau_{ij}\right)^{\sigma-1} \left(\frac{\sigma}{\sigma-1} a w_j\right)^{1-\sigma}\right)^{\frac{\mu}{1-\sigma}} \left(\frac{(1-\mu)}{H_i} \frac{w_i L_i}{\mu}\right)^{1-\mu}}\right)}{\sum_{i=1}^{\lambda} \exp\left(\dfrac{w_i/\chi}{\left(\sum_{j=1}^{\Lambda} \frac{L_j}{\sigma F} \left(\tau_{ij}\right)^{\sigma-1} \left(\frac{\sigma}{\sigma-1} a w_j\right)^{1-\sigma}\right)^{\frac{\mu}{1-\sigma}} \left(\frac{(1-\mu)}{H_i} \frac{w_i L_i}{\mu}\right)^{1-\mu}}\right)} = \frac{L_i}{L_A}, \text{for } i = 1, \ldots, \lambda-1.$$

$$\frac{\exp\left(\dfrac{w_i/\chi}{\left(\sum_{j=1}^{\Lambda} \frac{L_j}{\sigma F} \left(\tau_{ij}\right)^{\sigma-1} \left(\frac{\sigma}{\sigma-1} a w_j\right)^{1-\sigma}\right)^{\frac{\mu}{1-\sigma}} \left(\frac{(1-\mu)}{H_i} \frac{w_i L_i}{\mu}\right)^{1-\mu}}\right)}{\sum_{i=\lambda+1}^{\Lambda} \exp\left(\dfrac{w_i/\chi}{\left(\sum_{j=1}^{\Lambda} \frac{L_j}{\sigma F} \left(\tau_{ij}\right)^{\sigma-1} \left(\frac{\sigma}{\sigma-1} a w_j\right)^{1-\sigma}\right)^{\frac{\mu}{1-\sigma}} \left(\frac{(1-\mu)}{H_i} \frac{w_i L_i}{\mu}\right)^{1-\mu}}\right)} = \frac{L_i}{L_R}, \text{for } i = \lambda+2, \ldots, \Lambda.$$

Naturally,

$$\sum_{i=1}^{\lambda} L_i = L_A, \tag{22}$$

$$\sum_{i=\lambda+1}^{\Lambda} L_i = L_R, \tag{23}$$

where L_A, and L_R are the exogenously given country populations.

Overall equilibrium is characterized by the equilibrium values of 2Λ endogenous variables. These are the vector of nominal wages $[w_1, \ldots, w_\Lambda]$ and the vectors of the geographical distribution of labor in each country $[L_1, \ldots L_\lambda]$ and $[L_{\lambda+1}, \ldots L_\Lambda]$. We shall refer to this subset of endogenous variables as "core endogenous". The core endogenous variables are determined by the system of equations composed by the $\Lambda-1$ product–market equilibrium Eq. (19), the $\Lambda-2$ spatial equilibrium Eqs. (20), (21), and the two resource constraint Eqs. (22) and (23); which gives a total of $2\Lambda-1$ independent equations. We refer to this set of equations as the "core system". Choosing one endogenous variable as numéraire, the core system is perfectly determined. For notational convenience, we set $w_{\lambda+1} = 1$. Given the

[37] The parameters a, F, and the markup $\frac{\sigma}{\sigma-1}$ cancel out.

exogenous distribution of housing $\{H_i\}$ and the choice of numéraire, the core system determines the equilibrium vectors of the core endogenous variables: $[w_1^*,\ldots,w_\lambda^*,w_{\lambda+2}^*,\ldots,w_\Lambda^*]$, $[L_1^*,\ldots L_\lambda^*]$, and $[L_{\lambda+1}^*,\ldots L_\Lambda^*]$, where * denotes equilibrium values. Equilibrium values of all other endogenous variables can be computed from the equilibrium values of the core endogenous variables. Specifically, for each country the price of any variety obtains from expressions (13) and (14), the number of varieties obtains from expression (17), the price index obtains from Eq. (8), expenditure obtains from Eq. (10), the price of housing obtains from expression (9), and the real wage obtains from expression (11).

A.5. Three regions

For the purpose of our study, the model can be reduced to three regions, where A is composed of an interior region, I, and a border region, B, and R is a single-region country. Therefore, Eq. (23) and L_i for R drop out of the set of independent equations and from the set of endogenous variables, respectively. We are left with five $(2\Lambda-1)$ core endogenous variables: w_I, w_B, w_R, L_I, and L_B – of which we have already normalized $w_R=w_{\lambda+1}=1$ – and four $(2\Lambda-2)$ independent equations represented by the two Eq. in (19), the single equation in (20) and Eq. (22). It is useful to note that Eq. (20) may be rewritten as:

$$L_BP_I-L_IP_B=0. \tag{24}$$

The spatial equilibrium condition written in this way highlights the interpretation of the equilibrium as the state in which net migration flows are zero. Indeed, the first summand in Eq. (24) is the migration flow from region B to region I and the second summand is the migration flow from region I to region B. They must be equal in a spatial equilibrium. Writing Eq. (24) as $L_BP_I=L_IP_B$, taking the natural logarithm of both sides and rearranging gives:

$$\omega_B-\omega_I=\chi\ln\left(\frac{L_B}{L_I}\right). \tag{25}$$

In the numerical simulations, we therefore use the two Eq. (19), Eqs. (22), and (25), after having replaced the expression for real wages, to obtain w_I, w_B, L_I, and L_B.

Appendix B. Supplementary materials

Supplementary materials to this article can be found online at doi:10.1016/j.jinteco.2011.08.010.

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