How Wages and Employment Adjust to Trade Liberalization: Quasi-Experimental Evidence from Austria*

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Abstract

We study the responses 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 kilometers, we can identify statistically significant liberalization effects on both employment and wages. While wage responses preceded employment responses, the employment effect over the entire adjustment period is estimated to be three times as large as the wage effect. The implied slope of the regional labor supply curve can be replicated in a new economic geography model that features obstacles to labor migration due to immobile housing and to heterogeneous locational preferences.

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<All tables and figures at end>

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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 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, 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, so that market-access effects manifest themselves 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; 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.

The theoretical distinction between price and quantity effects of market access has been brought into focus by Head and Mayer (2004). Using a new economic 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, and we observe wage rises that precede the increases in employment.

We then seek to replicate our estimated magnitude of employment adjustment relative to wage adjustment in a calibrated three-region new economic geography model. A nontradeable 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 to this model a further dispersion force due to heterogeneous locational preferences, then we find that it predicts our central estimate of relative labor-market adjustment margins for realistic parameter calibrations.

The remainder of the paper is organized as follows. Section 2 describes the quasi-experimental empirical setting and the data. Our estimation strategy is described in Section 3, and results are reported in Section 4. In Section 5, we present a new economic geography model capable of reproducing our key estimated parameter. Section 6 concludes.

2 Empirical Setting and Data

2.1 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.

It should be uncontroversial to assume that the lifting of the Iron Curtain was exogenous to events in Austria. Moreover, over the period covered by our study, this transformation

took the form of a trade shock: a large change in cross-border openness of goods markets with little concomitant change in openness to cross-border worker flows.[1]

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.[2] 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.[3] Furthermore, the Eastern European countries all applied for full EU membership in the mid-1990s.[4] 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 emerges clearly from

[1] Free East-West mobility of workers has 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 labour 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.

[2] 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.

[3] Formally, these are the starting dates of "Interim Agreements". The official "Europe Agreeements" 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.

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

Figure 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 percent (Germany) and 20 percent (Switzerland), it increased by 107 percent with Hungary and by 178 percent with the Czech and Slovak republics. Figure 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 percent between 1992 and 2002. The data thus confirm that 1990 marked the start of a large and sustained eastward reorientation of Austrian trade.

Austria's small size implies that access to international markets is important: it was the OECD's fifth most trade oriented country in 1990.[5] Moreover, simple inspection of a map reveals that the transformations in Austria's eastern neighbors should have impacted differently on different regions of the country (see Figure 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.[6] Furthermore, we can draw on region-level data on foreign direct investment (FDI) collected by the Austrian central bank. In Figure 2, we report the share of the stock of outward FDI projects by Austrian firms that is located in Central and European Countries. It emerges clearly that firms in eastern Austria are significantly more oriented towards the eastern European markets than are 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 percent of total Austrian FDI in 1989 to 51 percent in 2002. In 2002, a full 96 percent of FDI from Austria's most easterly region (Burgenland) was invested in CEECs, while the corresponding share of Austria's most westerly

[5] According to the World Bank's World Development Indicators data base, Austria's trade-to-GDP ratio of 0.75 was surpassed only by Luxembourg, Belgium, Ireland and the Netherlands.

[6] See, for example, Combes, Lafourcade and Mayer (2005) and Helble (2007) for Europe, and Hillberry and Hummels (2003, 2008) for the United States.

region (Vorarlberg) was 23 percent. 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 percent, and the corresponding difference in effective wage rates amounted to fully 36 percent.[7] 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.

2.2 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.[8] These records can be matched to firms, 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 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 or municipality-sector.[9] Wages are recorded on a per-day basis,

[7] These data are taken from the 2002 statistical yearbook of the Austrian Federal Economic Chamber.

[8] 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.

[9] 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

which means that they are broadly comparable irrespective of whether employment contracts are part-time or full-time. A "day" in the ASSD corresponds to a calendar date, and not to a fully daily work period, which means that some of the variation in the wage data could nonetheless result from variation in hours worked. As a robustness test, we shall therefore consider data for males and females separately, in order to minimize potential distortions from differences in hours worked.

Firms are reported in 2,305 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.[10] Since we can allocate firms to one of 16 sectors, we can furthermore control for the industrial composition of municipalities.[11]

3 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:

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

where $\Delta Wage_{it}$ is the annual growth rate measured at quarterly intervals:

in all our sample years.

[10] 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.

[11] 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).

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

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.[12]

We then apply an equivalent specification for changes in municipal employment:

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

where $\Delta Empl$ is defined equivalently to $\Delta Wage$.

Our coefficients of interest are $\widehat{\alpha}$ and $\widehat{\beta}$. 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, $\rho = \frac{\beta}{\alpha}$, provides us with a measure of the relative magnitudes of nominal employment and wage adjustments, and thus of the slope of the average municipal labor supply curve. We shall take the estimated value of this ratio, $\widehat{\rho} = \frac{\widehat{\beta}}{\widehat{\alpha}}$, as the central empirical finding to be replicated by the theoretical model.[13]

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 estimate the following specifications:

$$\Delta Wage_{it} = \alpha_t(Border_i \times Fall_t \times d_t) + d_i + d_t + \nu_{it}^{wage}, \text{ and} \tag{3}$$

$$\Delta Empl_{it} = \beta_t(Border_i \times Fall_t \times d_t) + d_i + d_t + \nu_{it}^{empl}, \tag{4}$$

which gives us annual treatment effects $\widehat{\alpha}_t$ and $\widehat{\beta}_t$ for each year subsequent to the fall of the

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

[13] 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 amounts to estimating the slope of the regional labor supply curve, $\widehat{\rho}$, via indirect least squares.

Iron Curtain.

As a complement to parametric estimation, we in addition 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:

$$\Delta Wage_{it} = \gamma_i(Fall_t \times d_i) + d_i + d_t + \omega_{it}^{wage}, \text{ and} \tag{5}$$

$$\Delta Empl_{it} = \delta_i(Fall_t \times d_i) + d_i + d_t + \omega_{it}^{empl}. \tag{6}$$

The parameters $\widehat{\gamma}_i$ and $\widehat{\delta}_i$ 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).

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 a 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 alternatively using 3-sector and 16-sector classifications; and we then estimate the baseline specifications (1) and (2) for these matched samples.[14] In addition, we report 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, Duflo and Mullainathan, 2004).

[14] Our matching algorithm follows Abadie and Imbens (2006), allowing for multiple best matches (i.e. more than one matched interior municipality per border municipality in the case of "ties") and for matching with replacement (i.e. interior municipalities appear more than once if they represent a best match for more than one border municipality). The covariates (sectoral employment levels) are normalized by the inverse of their sample variances.

4 Results

4.1 Baseline Empirical Specification

For our baseline results, we define *Border* as comprising all municipalities whose geographic centre 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 Figure 3.

In Table 1, we present descriptive statistics separately for border and interior municipalities. The table shows that border municipalities had relatively low wages and small employment numbers in both subperiods. 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. We feel confident in assuming that no major shock coincided with that timing and geographic reach other than the opening of the Eastern markets.[15]

Our baseline econometric estimates are shown in Table 2. Column 1 reports the coefficient $\widehat{\alpha}$ from an estimation of the wage equation (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 treatment 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 $\widehat{\beta}$ from an estimation of equation (2), is given in column 3 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

[15] 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).

percent increase in nominal wages, and a 13 percent increase in employment, relative to regions in the Austrian interior.

Our estimated coefficients $\widehat{\alpha}$ and $\widehat{\beta}$ 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. $\widehat{\rho} = \frac{0.00861}{0.00267} = 3.2$).[16] Hence, employment was more responsive to changes in market access than nominal wages.

The estimated coefficients $\widehat{\alpha}$ and $\widehat{\beta}$ give us measures of the average effects of eastern market opening over the full 13-year treatment period covered by the data. We can get a description of the disaggregate time profile within that period by estimating specifications (3) and (4). These specifications provide us with annual estimates of differential wage changes ($\widehat{\alpha}_t$) and employment changes ($\widehat{\beta}_t$) in border regions for each year post-1990. The resulting estimates are shown in columns 2 and 4 of Table 2. In most sample years, border-region wage and employment growth rates did not statistically significantly diverge from those in interior regions. We do, however, observe two periods over which significant treatment effects are in evidence: nominal wages of border regions grew significantly more strongly in 1995-1997, and border-region employment growth showed a corresponding spike in 1997-2000. These periods are highlighted in Table 1. Our results suggest that wages adjust 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 but also to gradualism in the reduction of trade barriers and to persistence of policy risk (with fears of a political backlash in Eastern Europe persisting well into the 1990s).

In a nutshell, our baseline estimations imply that improved market access has boosted both employment and nominal wages in Austrian border regions, that the cumulative effect on employment was around three times bigger than the cumulative effect on wages, and that wages responded more quickly than employment levels.

[16] A Wald test shows that this effects is statistically significantly different from zero, but not from three, at the five-percent significance level.

4.2 Robustness

We subject our baseline results to a battery of robustness tests, reported in Tables 3 to 10. In all of these tables, we draw a box around the years that in the baseline estimation runs have appeared as the main periods of adjustment (i.e. 1995-1997 for wages and 1997-2000 for employment), so as to facilitate comparison.

4.2.1 Alternative Definitions of the Treatment Group

Our first exercise, reported in Table 3, is to consider also municipalities located between 25 and 50 kilometers from the eastern border. The coefficients 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 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-kilometre border zone. Experimentation with even wider border definitions never yielded any statistically significant results. A further result of our study, therefore, is that the differential market access effects were confined to a rather narrow set of locations in close proximity of the border.[17]

In a second robustness test, 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 minutes from a border crossing to the treatment sample.[18] The results, shown in Table 4, 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.[19]

[17] We provide further evidence of the steep spatial decay of the observed effects in Section 4.3.

[18] The overlap between the *Border* sample under the 25-kilometre 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.

[19] Figure 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 neighbour countries.

We report these results in Table 5. Qualitatively, our findings are equivalent to the baseline estimates. However, when dropping Slovenia, we find weaker evidence of a wage response and even stronger evidence of an employment response among the municipalities in the reduced-size treatment group. The average employment effect is now estimated to be 6.5 times as large as the average wage effect, and the latter is no longer statistically significant even at the ten-percent level. However, the baseline coefficient estimates are within the 95-percent confidence intervals of these alternative estimates. In terms of timing, we continue to observe employment adjustment occurring later than wage adjustment.

4.2.2 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 and itself situated close to the eastern border (see Figure 3). According to our distance measures, Vienna is located 64 kilometers, or 55 minutes, from the nearest eastern border (with Slovakia). It therefore is not included in our narrowly defined treatment groups. As it accounted for some 40 percent 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 Table 6, controlling for Vienna barely affects our baseline findings. However, we find a significant post-1990 boost to wages in Vienna. While there may have been unobserved factors at work that are specific to Vienna, some of the detected wage effect in Vienna could of course also be due to the renewed status of Vienna as a gateway to Eastern Europe.

One might furthermore suspect some of the measured market access effect to be due to the region of Burgenland. As shown in Figure 3, this region strongly overlaps with the set of municipalities defined as border regions with Hungary. Due to is relatively low per-capita income, Burgenland was granted Objective 1 status, making it eligible for generous regional subsidies, subsequent to Austria's accession to the European Union in 1995. 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 Table 6, columns 3-4 and 7-8. The Burgenland controls are negative in three of the four regression runs, but they are never statistically significant. Hence, Objective 1 status appears to have had no discernible impact on aggregate employment and wage growth in Burgenland. Most importantly, the inclusion of this control

variable also has no significant effect on our coefficient estimates of interest.

We next estimate our baseline models in samples of municipalities that are matched on industry-level employment shares. Thereby, we can examine whether our results are driven by the fact that border municipalities happened to be specialized in sectors that experienced particularly pronounced growth after 1990.[20] Columns 1-4 of Table 7 show the estimations for wage and employment growth with matching done over a coarse (3-sector) and over a finer (16-sector) industry classification. While the estimated wage effect approximately retains its magnitude, the estimated employment effect shrinks considerably and loses statistical significance. A plausible explanation for this can be found in Tables 3 and 6, which shows that the wage and employment effects do not fall off at equal rates relative to near-border (25-50 kilometer) and to Viennese municipalities. We therefore impose a second constraint on the control municipalities, by forcing them to be situated at least 70 kilometers from the eastern border, which just excludes Vienna. These results are shown in columns 5-8 of Table 7. While the standard errors are somewhat larger than in our baseline estimations, the point estimates are reassuringly close. Furthermore, the fact that the estimated employment effect exceeds the wage effect by a factor of three is perfectly consistent with the baseline results.

Much the same result appears in Table 8, where we report average treatment effects of a matching estimator applied to differences in growth rates between the post-1990 and the pre-1990 periods: again we find statistically significant treatment effects on employment as well as on wages, and a ratio of the two effects ($\widehat{\rho}$) of around three.[21]

4.2.3 Other Robustness Checks

As a further check on our baseline results, we estimate specifications (1) to (4) using weighted least squares regression and taking sample-average municipal employment as weights, so as to reduce the weight of very small municipalities. As shown in Table 9, this modification does not change our qualitative findings but increases the magnitudes and statistical significance of the relevant coefficients. The wage effect is now statistically significant at the one-percent level as well, with the employment effect estimated to be only 1.8 times as large as the wage

[20] 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 municipalities.

[21] Note that the coefficients reported in Table 8 cannot be directly compared to our baseline estimates, but to their implied cumulative effects of approximately 5 percent for wages and 13 percent for employment. Matching on industry composition shrinks these cumulative effects somewhat (to 3.3 and 10.1 percent respectively) without significantly affecting their ratio.

effect. The baseline coefficients, however, remain within the 95-percent confidence intervals also of these estimates. Employment responses still clearly lag the wage responses. While weighting is to some extent arbitrary, it is reassuring to see it strengthen the precision of our baseline estimates, as this measure lends greater weight to larger municipalities, for which it seems reasonable to assume that random fluctuations affecting a small number of workers play a relatively smaller role than for smaller municipalities.

We finally estimate our wage and employment regressions separately for male and female workers. Table 10 shows that this alters none of our conclusions. The ratio of wage to employment adjustment is somewhat higher for men ($\widehat{\rho}$ =3.2) than for women ($\widehat{\rho}$ =2.9), which is consistent with a more elastic female labor supply.[22]

4.3 Non-Parametric Illustration

So far, we have imposed a dichotomy between treatment ($Border = 1$) and control ($Border = 0$) municipalities. We now relax this by estimating specifications (5) and (6) and plotting the estimated post-1990 growth differential of each municipality against that municipality's distance from the eastern border. The plot for wages is given in Figure 4 and that for employment is given in Figure 6. Circles in these graphs are scaled according to municipal employment.

The raw scatter plots do not look particularly informative. The negative slope of a bivariate regression line fitted to those data is just about discernible. Nonetheless, a statistically significant relationship exists within the apparently featureless scatter. This becomes clear in the corresponding natural spline regressions shown in Figure 5 and Figure 7 respectively.[23] 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 kilometers from the border, with Vienna representing an evident outlier.

This representation therefore confirms that the differential effect of post-1990 market open-

[22] We have also run our estimations on industry-level data. All statistically significant effects of our treatment interaction variable are found to be positive also at the level of individual sectors, both for wages and for employment.

[23] 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.

ing was confined to a relatively narrow band of Austrian municipalities located close to the border. It would thus appear that in order to identify intra-national effects of trade liberalization, one has to zoom to a relatively small spatial scale.

5 Comparing Our Estimates to the Predictions of a New Economic Geography Model

We now compare our central estimates to the predictions implied by representative economic geography models. Our starting point is the variant of Krugman's (1991) "new economic geography" model proposed by Helpman (1998), which provides an attractive framework for the analysis of market-access effects at a relatively small spatial scale, as it explicitly considers congestion costs due to a non-tradeable resource H (thought of as housing).[24]

We then extend the Helpman model, by introducing heterogeneous locational preferences as an additional dispersion force, following Tabuchi and Thisse (2002) and Murata (2003).[25] As we will show, the extended model turns out to be more successful at reproducing our observed spatial adjustment patterns than the unamended Helpman model, the latter implying somewhat too elastic regional labor supply.

We provide details of the model in the Appendix and focus on sketching its main elements here. Labor, L, is the sole production factor, assumed to be perfectly mobile within a country. Worker-consumers spend a fraction μ of their income on varieties of a 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. The model features three regions, indexed by i: two regions in A(ustria) and one region in R(est of the world). A is composed of an interior region I and a border region B. Labor is fully employed and perfectly mobile within A but not between A and R. In the Helpman model, individuals decide where to locate according to the indirect utility they obtain from consumption of the tradeable good M and non-tradeable housing H. In the extended (Tabuchi-Thisse-Murata) model, we allow for randomly distributed idiosyncratic individual preferences for particular locations in addition to the purely pecuniary incentives considered in the Helpman model.

The non-linearity of the model makes it algebraically unsolvable. We therefore resort to

[24] Using this model will allow us to compare our results to those obtained by Redding and Sturm (2008).

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

numerical simulations.[26]

5.1 The Experiment

Our comparative-static exercise consists of tracking changes in nominal wages and employment within A as trade costs between A and R are lowered. Specifically, we seek a plausible and parsimonious model that predicts a value of the ratio of the relative change in employment to the relative change in the nominal wage that is close to the empirically observed value $\widehat{\rho}$.

Our baseline empirical results, reported in the first and third columns of Table 2, yield a $\widehat{\rho}$ of 3.2. The estimated average treatment effects when matching on industry compositions imply a $\widehat{\rho}$ of 3.0 (columns 1 and 2 of Table 8). Across our robustness tests, we have found this ratio to range between 1.8 to 6.5, with the majority of estimates in close vicinity of 3. This number lies within the 95-percent confidence intervals of all alternative estimates $\widehat{\rho}$. We thus take $\rho = 3$ as the ratio to be replicated.

Regions are separated by iceberg trade costs, whereby for each unit of merchandise 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 variety of the M-good to be transported from I to R it has to transit through B. Thus, the border region is nearer to R than the interior region.

We seek to model external trade liberalization of an integrated country. Therefore, 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. We choose the following values:

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

In order to generate a simulated value of ρ, we solve the model for each of the nine levels of τ_{BR}. Then, we compute the growth rate of nominal wages, $\widehat{w}_i$, and of employment, $\widehat{L}_i$,

[26] The Maple files used for the simulations are available from the authors.

for each 0.1 increment of trade cost reduction. It is important to note that these are growth rates between steady states. The empirical counterparts of these growth rates are the average or 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.

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

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

This ratio is observed for every increment of trade-cost reduction, which yields eight such ratios for each parameter combination. It turns out that the ratio varies only marginally across pairs of trade costs for which it is calculated. We therefore report averages of the eight computed ratios.

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.[27]

5.2 Parametrization

To calibrate the model, we need to decide on the values of the following parameters: housing stocks (in each region), population in A and R, the elasticity of substitution among differentiated goods, σ, the expenditure share of housing, $1-\mu$, and, in the extended model, the degree of heterogeneity in individuals' locational preferences. 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.[28]

As we shall see, the value assumed for $(1-\mu)$ 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 percent of the total wage bill

[27] 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 (see the upper-left parts of Tables 11 and 12). 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, demand and cost-of-living linkages, which favour instability, are attenuated compared to the original two-region model. Our extended model it is more stable still than the baseline model, since it contains an additional dispersion force in the form of taste heterogeneity.

[28] See, e.g., Baier and Bergstrand (2001), Bernard, Eaton, Jensen and Kortum (2003), Hanson (2005), Broda and Weinstein (2006) and Head and Mayer (2006).

and of 15 percent of the total wage bill plus net profits.[29]

The heterogeneity of locational preferences is modelled through the parameter $\chi \in (0, \infty)$.[30] When $\chi = 0$, individuals have identical preferences and choose their region of residence solely according to their indirect utility derived from the consumption of M and H. This is our starting point, as it corresponds to the preference structure of the Helpman (1998) model (and of most models of trade and geography). As χ increases, idiosyncratic locational preferences become more important, and in the extreme case of $\chi \to \infty$ they are all that matters for workers' location choices.

There is neither empirical nor theoretical guidance as to what value to assign to χ. We can, however, 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 then draw on some related empirical evidence, based on the mobility of unemployed workers (see Shields and Shields, 1989, for an early survey). Faini, Galli, Gennari and Rossi (1997) found that the percentage of Italian unemployed refusing to move out of their town of residence if a job were available elsewhere ranges from 20.7 percent (Northern male university graduates) to 61 percent (Southern low-education females). Fidrmuc (2005) reported survey evidence according to which 34 percent of EU15 unemployed and 24.9 percent 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.

Finally, we obtain housing stocks by calibrating the model such as to replicate the population distribution observed before the shock (for any combination of the other parameters).[31]

[29] Davis and Ortalo-Magné (2010) 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.

[30] See the Appendix for a precise definition.

[31] Municipalities we classify as belonging to the border region accounted for 5.1 percent of Austrian population pre-1990. Their implied housing stock in our calibrations ranges from 6 to 9 percent of the Austrian total.

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 simulations show that the implied ρs are almost unaffected by different parametrizations of H_R and L_R.[32]

5.3 Results

We proceed in two steps. First, we assume χ to be zero, as in the Helpman model, and we compute the predicted equilibrium values of ρ for the chosen parameter values on relative region sizes, substitution elasticities and expenditure shares of housing. These predicted values can be compared to our empirical estimate of around 3.

Second, we allow χ to take any non-negative value and search for the value of χ that yields an equilibrium ρ of 3. 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, which itself has no practical interpretation.[33]

The results of the first step are shown in Table 10, which reports the simulated values of ρ for several combinations of σ and $(1-\mu)$, with χ set to zero. As can be seen from the table, some of the simulations indeed yield values of ρ close to 3. However, for what could be considered our most realistic parameter combination, $\sigma = 4$ and $(1-\mu) = 0.25$, the predicted ρ, at 7.16, clearly exceeds the ratio suggested by our estimates.

Our simulations also show that the predicted value of ρ is not particularly sensitive to variations in σ but changes considerably with changes in $(1-\mu)$. In order to obtain a predicted 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 baseline model without heterogeneity of locational tastes predicts too much employment adjustment and too little wage adjustment. For a better

[32] Changes in ρ with respect to changes in these parameters only show up after the first decimal point. 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 attactiveness 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 - reflecting the difference-in-difference structure of the problem.

[33] 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 percent and the immobile population share were 95 percent, 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 component of the model.

match between the theory and our empirical result, a stronger dispersion force is needed than that represented by housing alone.

In the second step, we therefore allow for heterogeneity of locational preferences. The corresponding results are reported in Table 12. Each cell of that table shows the value of the preference parameter χ necessary for the implied equilibrium value of ρ to equal 3. Below these implied values of χ, we report the implied percentage real-wage differential between regions within country A (in parentheses) and the implied share of country A's population that prefers not to migrate at the prevailing real-wage differential (in brackets).

The results show 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 percent of the population would not move even if the real wage were 28 percent higher in the other region. In light of the available evidence on the issue, this does not appear to be an excessive dose of assumed intrinsic insensitivity to regional wage differentials.

5.4 Discussion

Our simulations suggest that the baseline new 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, this result stands in contrast to the findings of Hanson (2005) and Redding and Sturm (2008), who both concluded that the calibrated Helpman (1998) model fits 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 our three-region model and find levels of trade integration, τ_{BR}, for which the model precisely matches the estimated coefficient of the baseline employment regression, $\widehat{\beta} = 0.0086$ (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 excessive housing shares (of 42 and 48 percent 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 seems 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.[34] A comparison of his results to ours 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.

6 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 on the border to the formerly closed and centrally-planned economies and, on the other hand, Austrian regions further away from the border. We find that trade liberalization has had

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

statistically significant positive effects on both nominal wages and employment of the border regions, that these effects were confined to regions within less than 50 kilometers of the border, that wages adjusted faster than employment, and that the effect on employment exceeded the effect on nominal wages by a factor of around three.

We then calibrated a standard new economic geography model featuring immobile housing and compared the implied predictions to our estimation results. This comparison suggests that the model 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.

References

[1] Abadie, Alberto and Imbens, Guido (2006) Large Sample Properties of Matching Estimators for Average Treatment Effects. *Econometrica*, 74: 235-267.

[2] Ades, Alberto F. and Glaeser, Edward L. (1995) Trade and Circuses: Explaining Urban Giants. *Quarterly Journal of Economics*, 110: 195-227.

[3] Baier, Scott L. and Bergstrand, Jeffrey H. (2001) The Growth of World Trade: Tariffs, Transport Costs, and Income Similarity. *Journal of International Economics*, 53: 1-27.

[4] Baldwin, Richard E.; Forslid, Rikard; Martin, Philippe; Ottaviano, Gianmarco I.P. and Robert-Nicoud, Frédéric (2003) *Public Policies and Economic Geography.* Princeton University Press.

[5] Broda, Christian and Weinstein, David E. (2006) Globalization and the Gains from Variety. *Quarterly Journal of Economics*, 121: 541-585.

[6] Bernard, Andrew B.; Eaton, Jonathan; Jensen, J. Bradford and Kortum, Samuel (2003) Plants and Productivity in International Trade. *American Economic Review*, 93: 662-675.

[7] Bertrand, Marianne; Duflo, Esther and Mullainathan, Sendhil (2004) How Much Should We Trust Differences-in-Differences Estimates? *Quarterly Journal of Economics*, 119: 249-275.

[8] Buettner, Thiess and Rincke, Johannes (2007) Labor Market Effects of Economic Integration: The Impact of Re-Unification in German Border Regions. *German Economic Review*, 8: 536-560.

[9] Chiquiar, Daniel (2008) Globalization, Regional Wage Differentials and the Stolper-Samuelson Theorem: Evidence from Mexico. *Journal of International Economics*, 74: 70-93.

[10] Combes, Pierre-Philippe; Lafourcade, Miren and Mayer, Thierry (2005) The Trade-Creating Effects of Business and Social Networks: Evidence from France. *Journal of International Economics*, 66: 1-29.

[11] Davis, Donald R. and Weinstein, David E. (1999) Economic Geography and Regional Production Structure: An Empirical Investigation. *European Economic Review*, 43: 379-407.

[12] Davis, Morris A. and Ortalo-Magné, François (2010) Household Expenditures, Wages, Rents. *Review of Economic Dynamics*, forthcoming.

[13] Dickens, William T.; Goette, Lorenz; Groshen, Erica L.; Holden, Steinar; Messina, Julian; Schweitzer, Mark E.; Turunen, Jarkko and Ward, Melanie E. (2007) How Wages Change: Micro Evidence from the International Wage Flexibility Project. *Journal of Economic Perspectives*, 21: 195-214.

[14] Faber, Benjamin (2009) Integration and the Periphery: The Unintended Effects of New Highways in a Developing Country. *Mimeo*, London School of Economics.

[15] Faini, Riccardo; Galli, Gianpaolo; Gennari, Pietro and Rossi, Fulvio (1997) An Empirical Puzzle: Falling Migration and Growing Unemployment Differentials among Italian Regions. *European Economic Review*, 41: 571-579.

[16] Fidrmuc, Jan (2005) Labor Mobility During Transition: Evidence from the Czech Republic, *CEPR Discussion Paper* #5069.

[17] Hanson, Gordon H. (1997) Increasing Returns, Trade and the Regional Structure of Wages. *Economic Journal*, 107: 113-133.

[18] Hanson, Gordon H. (1998) Regional Adjustment to Trade Liberalization. *Regional Science and Urban Economics*, 28: 419-444.

[19] Hanson, Gordon H. (2005) Market Potential, Increasing Returns and Geographic Concentration. *Journal of International Economics*, 67: 1-24.

[20] Head, Keith and Mayer, Thierry (2004) The Empirics of Agglomeration and Trade. In: Henderson, J.V. and Thisse, J.F. (eds) *Handbook of Regional and Urban Economics*, Elsevier.

[21] Head, Keith, and Mayer, Thierry (2006) Regional Wage and Employment Responses to Market Potential in the EU. *Regional Science and Urban Economics*, 36: 573–594.

[22] Helble, Matthias (2007) Border Effect Estimates for France and Germany Combining International Trade and Intranational Transport Flows. *Review of World Economics*, 143: 433-463.

[23] Helpman, Elhanan (1998) The Size of Regions. In: Pines, D., Sadka, E. and Zilcha, I. (eds.) *Topics in Public Economics: Theoretical and Applied Analysis*, Cambridge University Press, 33-54.

[24] Henderson, J. Vernon (2003) The Urbanization Process and Economic Growth: The So-What Question. *Journal of Economic Growth*, 8: 47-71.

[25] Hillberry, Russell and Hummels, David (2003) Intranational Home Bias: Some Explanations. *Review of Economics and Statistics*, 85: 1089-1092.

[26] Hillberry, Russell and Hummels, David (2008) Trade Responses to Geographic Frictions: A Decomposition Using Micro-Data. *European Economic Review*, 52: 527-550.

[27] Krugman, Paul (1991) Increasing Returns and Economic Geography. *Journal of Political Economy*, 99: 483-499.

[28] Miyao, Takahiro (1978) A Probabilistic Model of Location Choice with Neighborhood Effects. *Journal of Economic Theory*, 19: 347–358.

[29] Murata, Yasusada (2003) Product Diversity, Taste Heterogeneity, and Geographic Distribution of Economic Activities: Market vs. Non-Market Interactions. *Journal of Urban Economics*, 53: 126-144.

[30] OECD (2001) *Migration Policies and EU Enlargement: The Case of Central and Eastern Europe*. OECD, Paris.

[31] Redding, Stephen and Sturm, Daniel (2008) The Costs of Remoteness: Evidence from German Division and Reunification. *American Economic Review*, 98: 1766-1797.

[32] Redding, Stephen and Venables, Anthony J. (2004) Economic Geography and International Inequality. *Journal of International Economics*, 62: 53-82.

[33] Sanguinetti, Pablo and Volpe Martincus, Christian (2009) Tariffs and Manufacturing Location in Argentina. *Regional Science and Urban Economics*, 39: 155-167.

[34] Shields, Gail M. and Shields, Michael P. (1989) The Emergence of Migration Theory and a Suggested New Direction. *Journal of Economic Surveys*, 3: 277-304.

[35] Tabuchi, Takatoshi and Thisse, Jacques-François (2002) Taste Heterogeneity, Labor Mobility and Economic Geography. *Journal of Development Economics*, 69: 155-177.

[36] Zweimüller, Josef; Winter-Ebmer, Rudolf; Lalive, Rafael; Kuhn, Andreas; Wuellrich, Jean-Philippe; Ruf, Oliver and Büchi, Simon (2009) Austrian Social Security Database. *Working Paper* #0903, Austrian Center for Labor Economics and the Welfare State, University of Linz.

A Appendix: 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).

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 from an exogenous and idiosyncratic preference parameter associated with individual regions.

The component of utility that is associated with consumption is modelled 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:

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

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:

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

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

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

In our extended model (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

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

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 $\lim_{\chi\to\infty} \Pr_i\left(\omega_i, \chi\right) = \frac{1}{\lambda}$ (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, $\lim_{\chi\to 0} \Pr_i\left(\omega_i, \chi\right) = \frac{1}{\sum_{j=1}^{\lambda} \exp(\omega_j/\omega_i)}$, 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 $\lim_{\chi\to\infty} \Pr_i\left(\omega_i, \chi\right) = \frac{1}{\Lambda-\lambda}$ and $\lim_{\chi\to 0} \Pr_i\left(\omega_i, \chi\right) = \frac{1}{\sum_{j=\lambda+1}^{\Lambda} \exp(\omega_j/\omega_i)}$.

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 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:

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

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

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 equation (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:

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

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 (equations (13)-(14)), the price index (equation (8)), total expenditure (equation (10)), the number of varieties (equation (17)), and equilibrium output of any variety (equation (15)) into (18), the system of market-clearing equations for a given variety of M becomes:[35]

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

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

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, 2003). Thus, a spatial equilibrium requires the following:

$$L_A \Pr{}_i(\omega_i, \chi) = L_i, \quad \text{for } i = 1, ..\lambda - 1, \tag{20}$$

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

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

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

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

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

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 $[w_1, ..., w_\Lambda]$ and the vectors of the geographical distribution of labor in each country $[L_1, ...L_\lambda]$ and $[L_{\lambda+1}, ...L_\Lambda]$. 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 equations (19), the $\Lambda - 2$ spatial equilibrium equations (20), (21), and the two resource constraint equations (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 exogenous distribution of housing $\{H_i\}$ and the choice of numéraire, the core system determines the equilibrium vectors of the core endogenous variables: $\left[w_1^*, ..., w_\lambda^*, w_{\lambda+2}^*, ..., w_\Lambda^*\right]$, $[L_1^*, ...L_\lambda^*]$, and $\left[L_{\lambda+1}^*, ...L_\Lambda^*\right]$, where * denotes equilibrium values. Equilibrium values of all other endoge-

nous 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 (8), expenditure obtains from (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, equation (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 equations in (19), the single equation in (20) and equation (22). It is useful to note that equation (20) may be rewritten as:

$$L_B P_I - L_I P_B = 0. \tag{24}$$

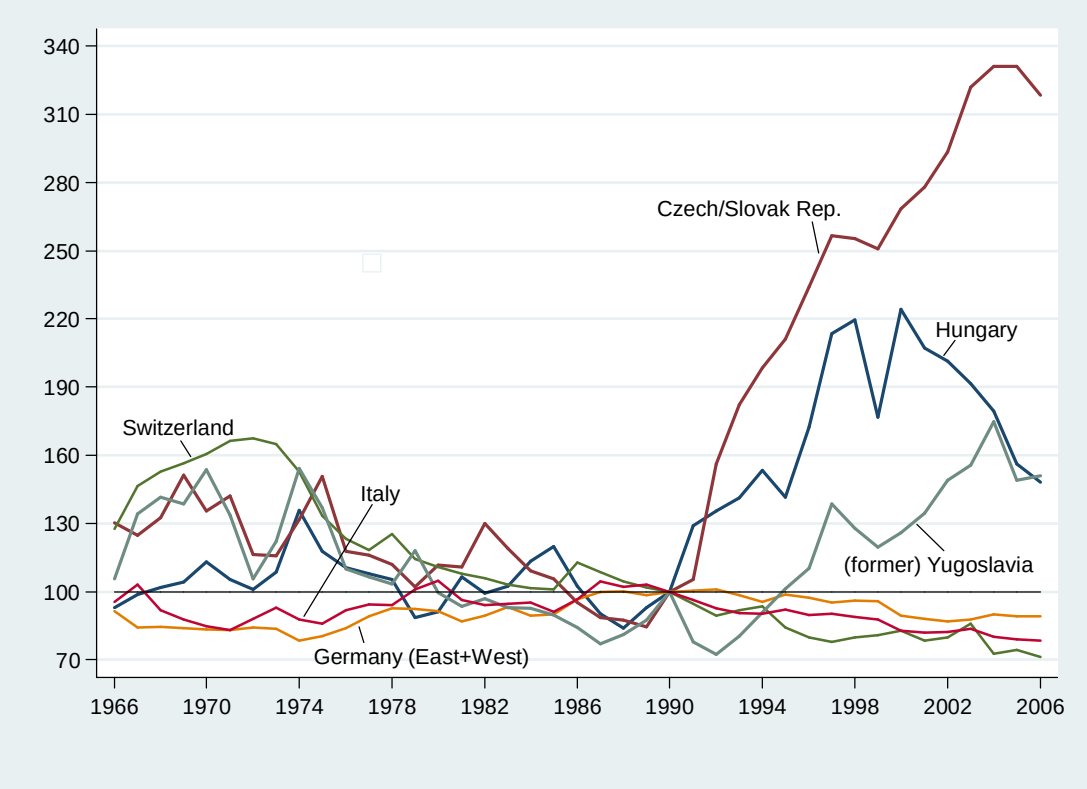
The spatial equilibrium condition written in this way highlights the interpretation of the equilibrium as the state in which net migration flows are zero. Indeed, the first summand in equation (24) is the migration flow from region B to region I and the second summand is the migration flow from region I to region B. They must be equal in a spatial equilibrium. Writing (24) as $L_B P_I = L_I P_B$, 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 equations (19), equation (22), and equation (25), after having replaced the expression for real wages, to obtain w_I, w_B, L_I, and L_B.

Figure 1: Austria's Post-1990 Eastward Trade Opening

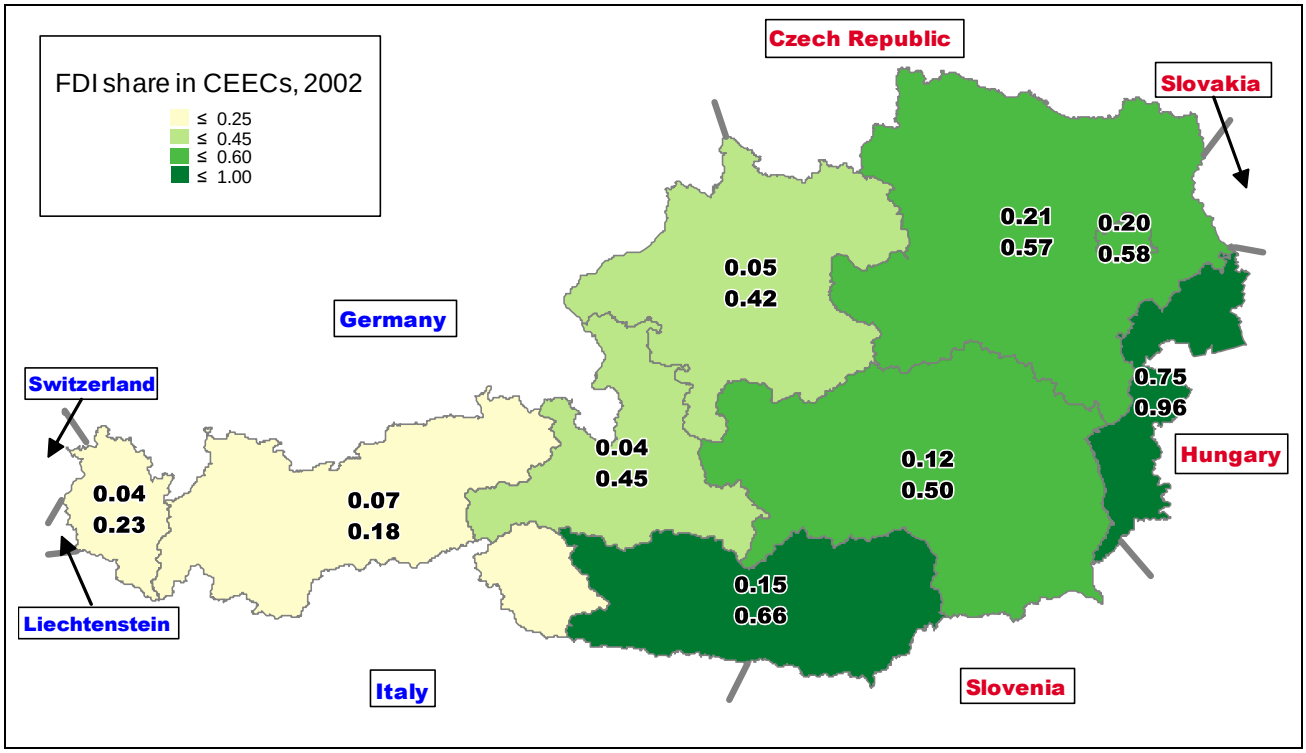
Value of merchandise imports + exports as share of total Austrian trade, 1990 = 100



Source: UN Comtrade database

Figure 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

Figure 3: Treatment Municipalities

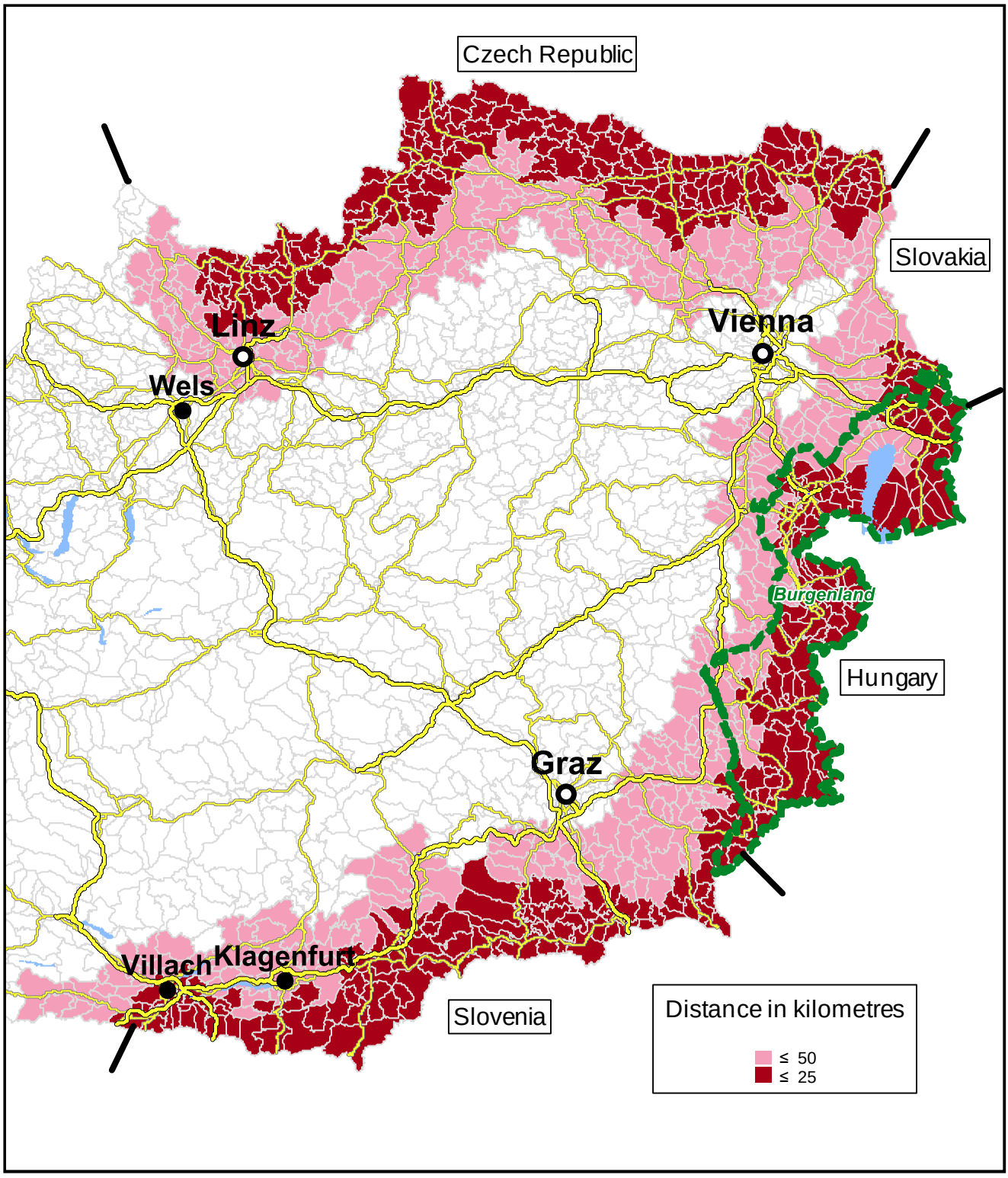


Figure 4: Distance to Border and Post-1990 Wage Growth

Each point/circle represents a municipality; circle sizes are proportional to total employment; vertical axis represents differential post-1990 growth in median nominal wages ($\hat{\gamma}_i$ of equation 5)

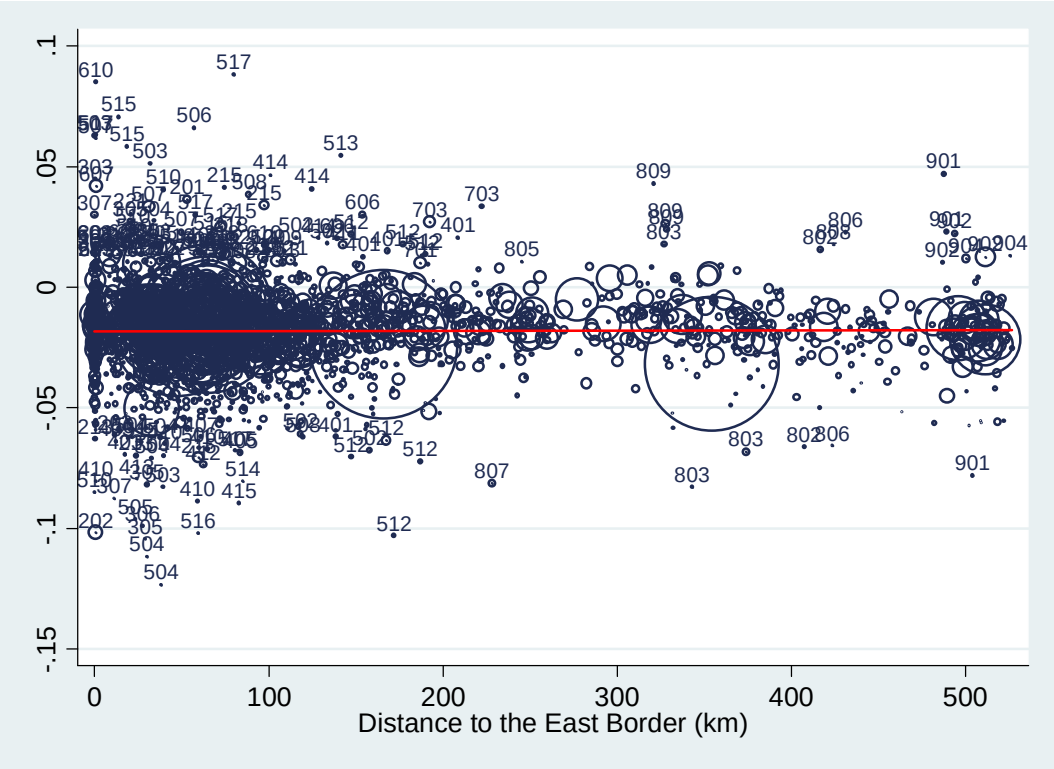


Figure 5: Distance to Border and Post-1990 Wage Growth – Nonparametric Fit

Natural spline regression on data shown in Figure 4 (7 degrees of freedom); 95% confidence interval

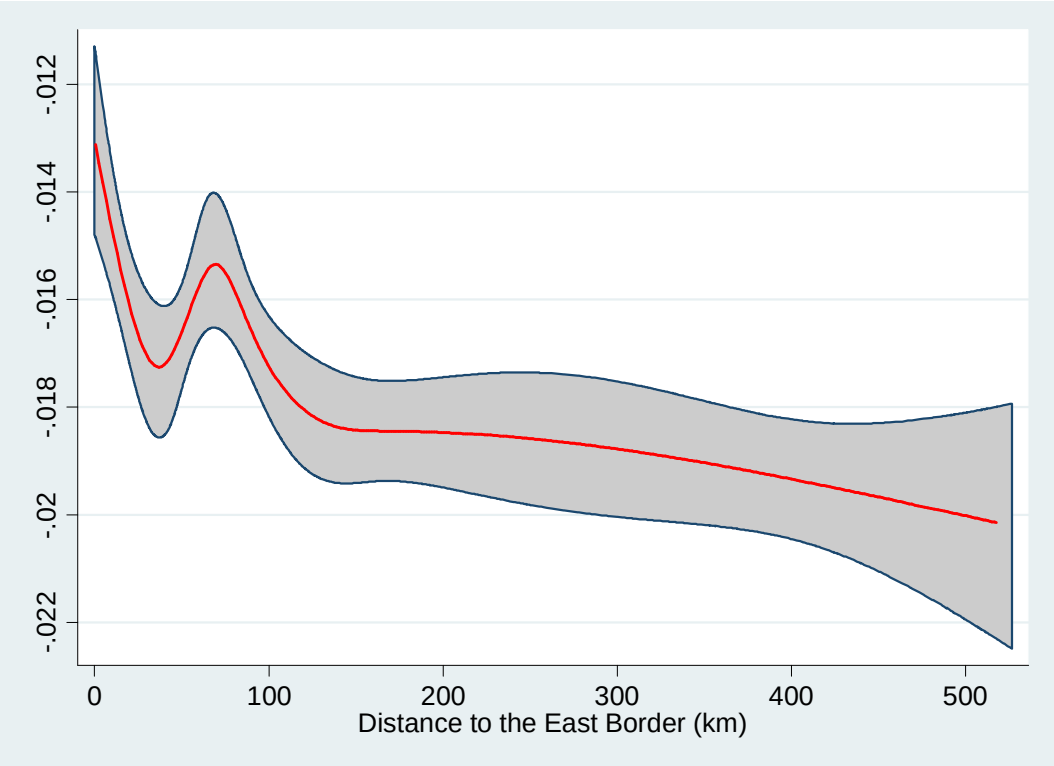


Figure 6: Distance to Border and Post-1990 Employment Growth

Each point/circle represents a municipality; circle sizes are proportional to total employment; vertical axis represents differential post-1990 growth in total employment ($\hat{\delta}_i$ of equation 6)

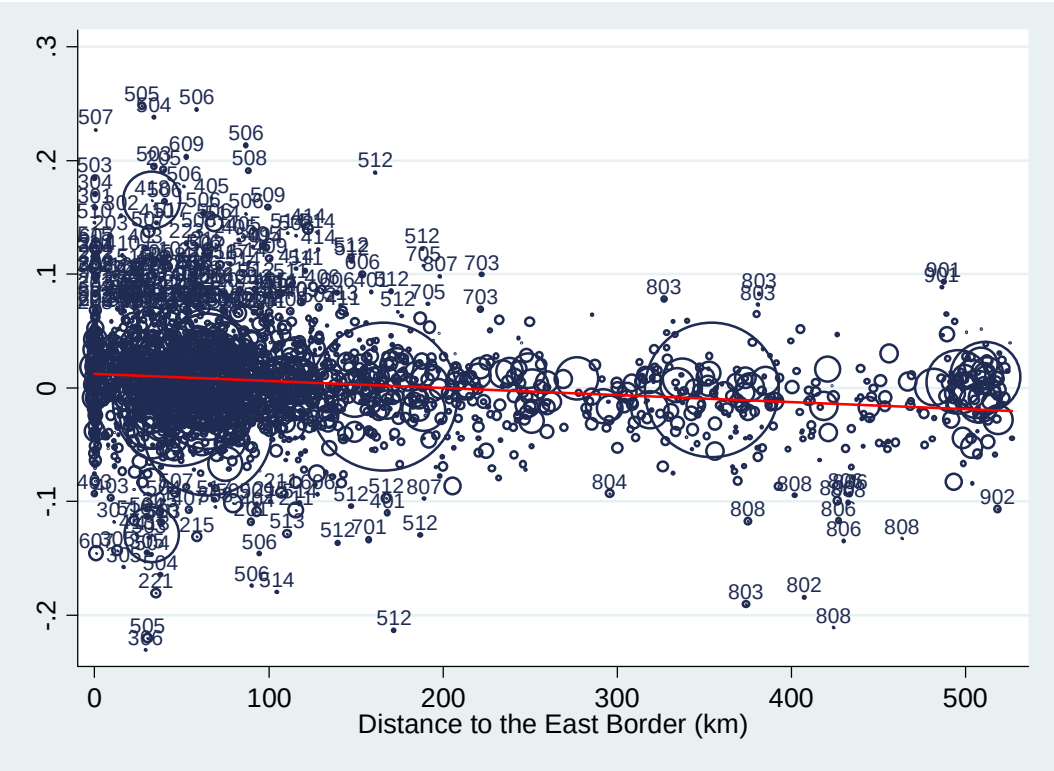


Figure 7: Distance to Border and Post-1990 Employment Growth – Nonparametric Fit

Natural spline regression on data shown in Figure 6 (7 degrees of freedom); 95% confidence intervals

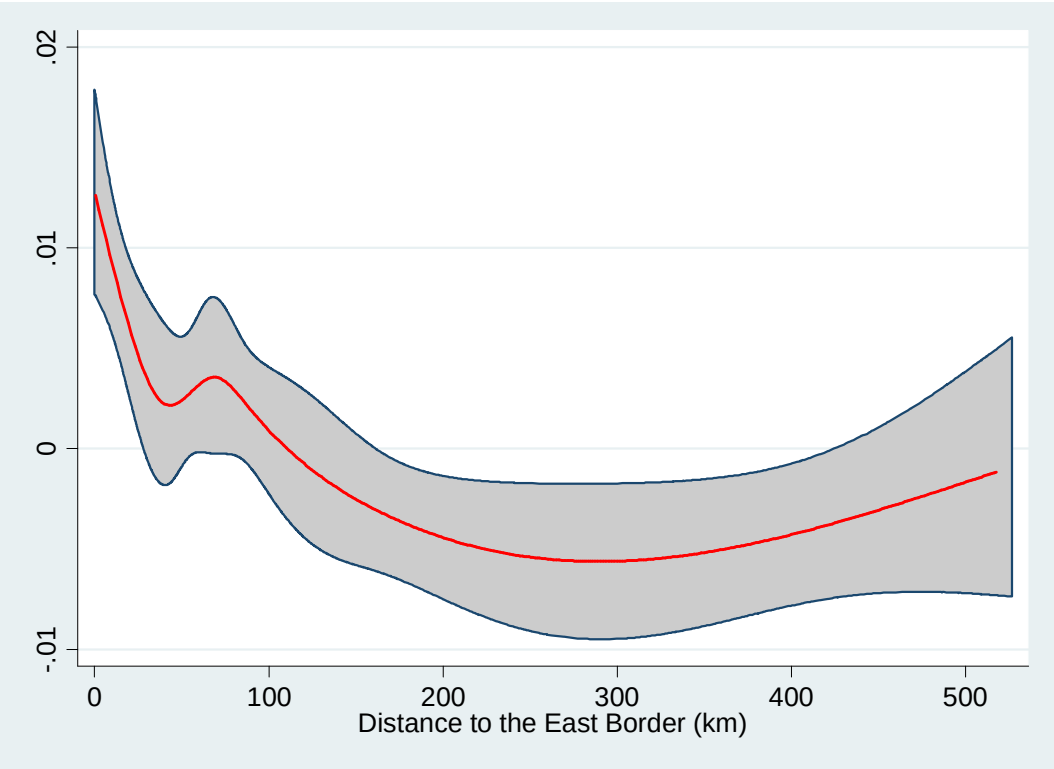


Table 1: Summary Statistics

Border defined as municipalities within 25 km from Czech, Hungarian, Slovakian or Slovenian border

	1975-1989				1990-2002			
Variables	**Mean**	**Std. Dev.**	**Min**	**Max**	**Mean**	**Std. Dev.**	**Min**	**Max**
Border municipalities (*Border* = 1)	18,760 observations: 56 quarters and 335 municipalities				17,420 observations: 52 quarters and 335 municipalities			
Median daily wage (Austrian Schillings)	321.22	97.20	53.10	797.44	626.16	139.31	191.26	1,660.97
Annual growth rate of median wage, $\Delta Wage$	0.0616	0.09	-1.03	1.28	0.0413	0.09	-1.01	0.87
Employment	307.65	863.95	1	15,843	351.43	974.83	1	20,185
Annual growth rate of employment, $\Delta Empl$	0.0113	0.14	-1.63	1.65	0.0179	0.22	-1.85	1.86
Minimum road distance to Eastern border (kilometres)	10.83	8.88	0.00	24.99	10.83	8.88	0.00	24.99
Minimum road travelling time to Eastern border (minutes)	20.47	10.29	0.00	48.17	20.47	10.29	0.00	48.17
Interior municipalities (*Border* = 0)	110,320 observations: 56 quarters and 1,970 municipalities				102,440 observations: 52 quarters and 1,970 municipalities			
Median daily wage (Austrian Schillings)	354.90	109.50	35.52	950.62	651.73	145.95	119.72	1,578.62
Annual growth rate of median wage, $\Delta Wage$	0.0585	0.09	-1.35	1.51	0.0356	0.09	-1.21	1.25
Employment	968.77	4,319.09	1	87,379	1,023.34	4,524.56	1	94,095
Annual growth rate of employment, $\Delta Empl$	0.0133	0.14	-1.86	1.82	0.0112	0.19	-1.83	1.84
Minimum road distance to Eastern border (kilometres)	139.74	132.22	25.00	526.56	139.74	132.22	25.00	526.56
Minimum road travelling time to Eastern border (minutes)	117.80	79.63	29.00	375.33	117.80	79.63	29.00	375.33

Table 2: Baseline Regressions

Border defined as municipalities within 25 km from Czech, Hungarian, Slovakian or Slovenian border

	Dependent variable:			
	ΔWage		***ΔEmpl***	
	(1)	(2)	(3)	(4)
Border x Fall, 1990-2002	**0.00267****		**0.00861*****	
	(0.00121)		(0.00281)	
Border x Fall, 1990		0.00481		0.0119*
		(0.00366)		(0.00670)
Border x Fall, 1991		-0.00126		0.00193
		(0.00370)		(0.00737)
Border x Fall, 1992		0.00560*		0.00254
		(0.00337)		(0.00800)
Border x Fall, 1993		0.00150		0.00472
		(0.00341)		(0.00652)
Border x Fall, 1994		0.00148		0.000123
		(0.00406)		(0.00802)
Border x Fall, 1995		0.00671*		-0.00564
		(0.00375)		(0.00647)
Border x Fall, 1996		0.0120***		0.00265
		(0.00395)		(0.00738)
Border x Fall, 1997		0.00642**		0.00701*
		(0.00320)		(0.00446)
Border x Fall, 1998		-0.00199		0.00504
		(0.00682)		(0.0197)
Border x Fall, 1999		-0.00588		0.0811***
		(0.00709)		(0.0194)
Border x Fall, 2000		-0.000368		0.0121**
		(0.00345)		(0.00604)
Border x Fall, 2001		0.00326		-0.00461
		(0.00341)		(0.00637)
Border x Fall, 2002		0.00238		0.00316
		(0.00393)		(0.00917)
No. obs.	248,940	248,940	248,940	248,940
No. municipalities	2,305	2,305	2,305	2,305
R^2	0.058	0.058	0.176	0.177
quarter fixed effects	Yes	Yes	Yes	Yes
municip. fixed effects	Yes	Yes	Yes	Yes

Note: estimation with OLS; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; * : p=0.1, **: p=0.05, ***: p=0.01

Table 3: Wider Definition of "Border" Municipalities

Border defined as Municipalities within 25 or within 25-50 km from Czech, Hungarian, Slovakian or Slovenian border

	Dependent variable:			
	ΔWage		***ΔEmpl***	
	(1)	(2)	(3)	(4)
Border 0-25 x Fall, 1990-2002	**0.00231***		**0.00994*****	
	(0.00122)		(0.00286)	
Border 25-50 x Fall, 1990-2002	**0.00143**		**0.00533****	
	(0.00105)		(0.00265)	
Border 0-25 x Fall, 1990		0.00503		0.0120*
		(0.00371)		(0.00671)
Border 0-25 x Fall, 1991		-0.00104		0.00135
		(0.00377)		(0.00748)
Border 0-25 x Fall, 1992		0.00427		0.00206
		(0.00339)		(0.00806)
Border 0-25 x Fall, 1993		0.00153		0.00866
		(0.00346)		(0.00662)
Border 0-25 x Fall, 1994		0.00073		0.000915
		(0.00413)		(0.00812)
Border 0-25 x Fall, 1995		0.00646*		-0.00401
		(0.00379)		(0.00660)
Border 0-25 x Fall, 1996		0.0114***		0.00212
		(0.00399)		(0.00744)
Border 0-25 x Fall, 1997		0.00664**		0.00834
		(0.00322)		(0.00554)
Border 0-25 x Fall, 1998		0.000343		-0.00151
		(0.00698)		(0.02000)
Border 0-25 x Fall, 1999		-0.0106		0.0828***
		(0.00723)		(0.01970)
Border 0-25 x Fall, 2000		-0.000192		0.0143**
		(0.00353)		-0.00614
Border 0-25 x Fall, 2001		0.00313		-0.00337
		(0.00347)		-0.00651
Border 0-25 x Fall, 2002		0.00231		0.00554
		(0.00404)		(0.00936)
Border 25-50 x Fall, 1990		0.000856		0.000463
		(0.00411)		(0.00776)
Border 25-50 x Fall, 1991		0.000872		-0.00232
		(0.00342)		(0.00635)
Border 25-50 x Fall, 1992		-0.00534		-0.00194
		(0.00344)		(0.00668)
Border 25-50 x Fall, 1993		0.000131		0.0158***
		(0.00287)		(0.00567)
Border 25-50 x Fall, 1994		-0.003		0.00317
		(0.00314)		(0.00656)
Border 25-50 x Fall, 1995		-0.000964		0.00652
		(0.00282)		(0.00578)
Border 25-50 x Fall, 1996		0.00262		-0.0021
		(0.00316)		(0.00575)
Border 25-50 x Fall, 1997		0.000895		0.00534
		(0.00326)		(0.00578)
Border 25-50 x Fall, 1998		0.00933**		0.0141
		(0.00469)		(0.01470)
Border 25-50 x Fall, 1999		-0.0187***		0.00688
		(0.00498)		(0.01460)
Border 25-50 x Fall, 2000		0.000704		0.00891*
		(0.00336)		(0.00496)
Border 25-50 x Fall, 2001		-0.000502		0.00495
		(0.00329)		(0.00520)
Border 25-50 x Fall, 2002		-0.000281		0.00953
		(0.00407)		(0.00798)
No. obs.	248,940	248,940	248,940	248,940
No. municipalities	2,305	2,305	2,305	2,305
R^2	0.058	0.058	0.176	0.177
quarter fixed effects	Yes	Yes	Yes	Yes
municip. fixed effects	Yes	Yes	Yes	Yes

Note: estimation with OLS; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; * : p=0.1, **: p=0.05, ***: p=0.01

Table 4: Definition of "Border" Municipalities in Terms of Travel Time

Border defined as municipalities within 35 minutes' traveling time from nearest Czech, Hungarian, Slovakian or Slovenian border crossing

	Dependent variable:			
	ΔWage		***ΔEmpl***	
	(1)	(2)	(3)	(4)
Border x Fall, 1990-2002	**0.00271****		**0.00877****	
	(0.00131)		(0.00408)	
Border x Fall, 1990		0.00331		0.0179**
		(0.00437)		(0.00779)
Border x Fall, 1991		0.000992		0.00255
		(0.00408)		(0.00835)
Border x Fall, 1992		0.00282		0.000518
		(0.00372)		(0.00847)
Border x Fall, 1993		0.000515		0.00750
		(0.00379)		(0.00732)
Border x Fall, 1994		0.00144		-0.00374
		(0.00459)		(0.00922)
Border x Fall, 1995		0.00756*		0.000850
		(0.00420)		(0.00715)
Border x Fall, 1996		0.00893**		0.0151*
		(0.00418)		(0.00851)
Border x Fall, 1997		0.00967***		0.00889
		(0.00311)		(0.00563)
Border x Fall, 1998		-0.000308		0.0228
		(0.00662)		(0.0200)
Border x Fall, 1999		-0.00545		0.0343*
		(0.00673)		(0.0197)
Border x Fall, 2000		-0.000127		0.0166***
		(0.00365)		(0.00611)
Border x Fall, 2001		0.00243		-0.00763
		(0.00389)		(0.00708)
Border x Fall, 2002		0.00347		-0.00162
		(0.00417)		(0.00996)
No. obs.	248,940	248,940	248,940	248,940
No. municipalities	2,305	2,305	2,305	2,305
R^2	0.058	0.058	0.162	0.162
quarter fixed effects	Yes	Yes	Yes	Yes
municip. fixed effects	Yes	Yes	Yes	Yes

Note: estimation with OLS; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; * : p=0.1, **: p=0.05, ***: p=0.01

Table 5: Definition of "Border" Municipalities Without Slovenia

Border defined as municipalities within 25 km from Czech, Hungarian or Slovakian border

	Dependent variable:			
	ΔWage		***ΔEmpl***	
	(1)	(2)	(3)	(4)
Border x Fall, 1990-2002	**0.00150**		**0.00973*** **	
	(0.00136)		(0.00332)	
Border x Fall, 1990		0.000884		0.0179**
		(0.00407)		(0.00749)
Border x Fall, 1991		-0.00287		0.00456
		(0.00438)		(0.00888)
Border x Fall, 1992		0.00399		-0.00234
		(0.00410)		(0.00997)
Border x Fall, 1993		0.00119		0.00331
		(0.00363)		(0.00762)
Border x Fall, 1994		0.00460		-0.000294
		(0.00436)		(0.0100)
Border x Fall, 1995		0.00530*		-0.0106
		(0.00326)		(0.00790)
Border x Fall, 1996		0.00714*		0.000179
		(0.00444)		(0.00839)
Border x Fall, 1997		0.00305		0.00950
		(0.00355)		(0.00613)
Border x Fall, 1998		-0.00447		0.0552**
		(0.00882)		(0.0237)
Border x Fall, 1999		-0.00483		0.111***
		(0.00960)		(0.0243)
Border x Fall, 2000		0.00233		0.0172**
		(0.00435)		(0.00737)
Border x Fall, 2001		0.00104		-0.000140
		(0.00384)		(0.00724)
Border x Fall, 2002		0.00210		0.0320***
		(0.00463)		(0.0103)
No. obs.	248,940	248,940	248,940	248,940
No. municipalities	2,305	2,305	2,305	2,305
R^2	0.058	0.058	0.176	0.178
quarter fixed effects	Yes	Yes	Yes	Yes
municip. fixed effects	Yes	Yes	Yes	Yes

Note: estimation with OLS; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; * : p=0.1, **: p=0.05, ***: p=0.01

Table 6: Controlling for Vienna and Burgenland

Border defined as municipalities within 25 km from Czech, Hungarian, Slovakian or Slovenian border

	Dependent variable:							
	ΔWage				***ΔEmpl***			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Border x Fall, 1990-2002	**0.00272****		**0.00287****		**0.00862*****		**0.00837*****	
	(0.00121)		(0.0013)		(0.00281)		(0.003)	
Vienna dummy x Fall, 1990-2002	**0.00405*****	0.00405***	**0.00404*****	0.00404***	**0.00128**	0.00128	**0.00129**	0.00126
	(0.00103)	(0.00103)	(0.001)	(0.0010)	(0.00351)	(0.00351)	(0.0035)	(0.0035)
Burgenland dummy x 95-02 dummy			**-0.000929**	-0.00118			**0.00157**	-0.00104
			(0.0021)	(0.0022)			(0.0057)	(0.0061)
Border x Fall, 1990		0.00486		0.00486		0.0119*		0.0119*
		(0.00366)		(0.0037)		(0.00670)		(0.0067)
Border x Fall, 1991		-0.00121		-0.00121		0.00194		0.00194
		(0.00370)		(0.0037)		(0.00738)		(0.0074)
Border x Fall, 1992		0.00565*		0.00565*		0.00256		0.00256
		(0.00337)		(0.0034)		(0.00800)		(0.0080)
Border x Fall, 1993		0.00154		0.00154		0.00474		0.00474
		(0.00341)		(0.0034)		(0.00652)		(0.0065)
Border x Fall, 1994		0.00153		0.00153		0.000138		0.000138
		(0.00406)		(0.0041)		(0.00802)		(0.0080)
Border x Fall, 1995		0.00675*		0.00707*		-0.00563		-0.00535
		(0.00375)		(0.0038)		(0.00647)		(0.0068)
Border x Fall, 1996		0.0121***		0.0124***		0.00266		0.00294
		(0.00395)		(0.0039)		(0.00738)		(0.0076)
Border x Fall, 1997		0.00647**		0.00678**		0.00702		0.00730
		(0.00320)		(0.0032)		(0.00546)		(0.0056)
Border x Fall, 1998		-0.00194		-0.00163		0.00502		-0.00475
		(0.00682)		(0.0068)		(0.0197)		(0.020)
Border x Fall, 1999		-0.00583		-0.00552		0.0811***		0.0814***
		(0.00709)		(0.0072)		(0.0194)		(0.020)
Border x Fall, 2000		-0.000321		-0.00001		0.0121**		0.0123**
		(0.00345)		(0.0035)		(0.00604)		(0.0062)
Border x Fall, 2001		0.00330		0.00362		-0.00459		-0.00432
		(0.00341)		(0.0035)		(0.00637)		(0.0065)
Border x Fall, 2002		0.00243		0.00274		0.00317		0.00345
		(0.00393)		(0.0040)		(0.00917)		(0.0094)
No. obs.	248,940	248,940	248,940	248,940	248,940	248,940	248,940	248,940
No. municipalities	2,305	2,305	2,305	2,305	2,305	2,305	2,305	2,305
R^2	0.058	0.058	0.058	0.058	0.176	0.177	0.177	0.177
quarter fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
municip. fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: estimation with OLS; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; * : p=0.1, **: p=0.05, ***: p=0.01

Table 7: Controlling for Industrial Composition: Regression

Sample of interior municipalities confined to those that provide best match to one or several border municipalities in terms of sector-level employment

	Dependent variable:							
	ΔWage	***ΔEmpl***	***ΔWage***	***ΔEmpl***	***ΔWage***	***ΔEmpl***	***ΔWage***	***ΔEmpl***
	(1)	(2)	(4)	(5)	(5)	(6)	(7)	(8)
Border x Fall, 1990-2002	**0.00224****	**0.00311**	**0.00387****	**0.00638**	**0.00409****	**0.00932****	**0.00282***	**0.00846****
	(0.00135)	(0.00410)	(0.00162)	(0.00424)	(0.00164)	(0.00387)	(0.00165)	(0.00415)
Matching on	1989 employment in 3 sectors		1989 employment in 16 sectors		1989 employment in 3 sectors *and* geographic constraint		1989 employment in 16 sectors *and* geographic constraint	
No. obs.	71,496	71,496	67,824	67,824	67,608	67,608	62,100	62,100
No. municipalities	662	662	628	628	626	626	575	575
R^2	0.049	0.163	0.051	0.166	0.050	0.183	0.052	0.179
quarter fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
municip. fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: estimation with OLS; covariate matching with replacement by minimising the sum of squared differences in normalised 1989 sector-level employment; geographic constraint: control group (non-border muncipalities) ≥70km from eastern border; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; * : p=0.1, **: p=0.05, ***: p=0.01

Table 8: Controlling for Industrial Composition: Matching

Average treatment effects: cumulative growth differential post-minus-pre-1990, border versus interior municipalities

	Dependent variable:			
	ΔWage_cumul	***ΔEmpl_cumul***	***ΔWage_cumul***	***ΔEmpl_cumul***
	(1)	(2)	(3)	(4)
Border	**0.0330***	**0.1006****	**0.0337***	**0.1174*****
	(0.01690)	(0.03920)	(0.01760)	(0.04080)
Matching on	1989 employment in 16 sectors		1989 employment in 16 sectors *and* geographic constraint	
No. obs.	2,305		1,573	

Note: dependent variables defined as difference between cumulative growth post-1990 and cumulative growth pre-1990; covariate matching with replacement by minimising the sum of squared differences in normalised 1989 sector-level employment, four nearest neighbours; geographic constraint: control group (non-border muncipalities) ≥70km from eastern border; standard errors in parentheses; * : p=0.1, **: p=0.05, ***: p=0.01

Table 9: Weighted Least Squares

Border defined as municipalities within 25 km from Czech, Hungarian, Slovakian or Slovenian border

	Dependent variable:			
	ΔWage		***ΔEmpl***	
	(1)	(2)	(3)	(4)
Border x Fall, 1990-2002	**0.00458*** **		**0.00815*** **	
	(0.00120)		(0.00313)	
Border x Fall, 1990		0.00116		-0.00131
		(0.00263)		(0.00688)
Border x Fall, 1991		0.00163		-0.00305
		(0.00221)		(0.00449)
Border x Fall, 1992		0.00504		0.00281
		(0.00364)		(0.00621)
Border x Fall, 1993		0.00443**		0.00562
		(0.00223)		(0.00575)
Border x Fall, 1994		0.00550***		0.0111**
		(0.00211)		(0.00564)
Border x Fall, 1995		0.0112***		0.0110*
		(0.00267)		(0.00635)
Border x Fall, 1996		0.00855***		0.0129**
		(0.00264)		(0.00540)
Border x Fall, 1997		0.00946***		0.0167***
		(0.00241)		(0.00517)
Border x Fall, 1998		0.00135		-0.0853
		(0.0116)		(0.0560)
Border x Fall, 1999		0.000894		0.157***
		(0.0103)		(0.0546)
Border x Fall, 2000		0.0000860		0.00517
		(0.00266)		(0.00483)
Border x Fall, 2001		0.00881**		0.000921
		(0.00353)		(0.00650)
Border x Fall, 2002		0.00150		-0.0277*
		(0.00281)		(0.0144)
No. obs.	248,940	248,940	248,940	248,940
No. municipalities	2,305	2,305	2,305	2,305
R^2	0.269	0.269	0.294	0.299
quarter fixed effects	Yes	Yes	Yes	Yes
municip. fixed effects	Yes	Yes	Yes	Yes

Note: estimation with weighted least squares (population weights); standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; * : p=0.1, **: p=0.05, ***: p=0.01

Table 10: Gender Decomposition

Border defined as municipalities within 25 km from Czech, Hungarian, Slovakian or Slovenian border

	Men only		**Women only**	
	Dependent variable:		*Dependent variable:*	
	ΔWage	***ΔEmpl***	***ΔWage***	***ΔEmpl***
	(1)	(2)	(3)	(4)
Border x Fall, 1990-2002	**0.00245****	**0.00787****	**0.00334****	**0.00969****
	(0.00120)	(0.00342)	(0.00151)	(0.00427)
No. obs.	248,203	248,586	247,931	248,456
No. municipalities	2,305	2,305	2,305	2,305
R^2	0.045	0.11	0.052	0.139
quarter fixed effects	Yes	Yes	Yes	Yes
municip. fixed effects	Yes	Yes	Yes	Yes

Note: estimation with OLS; standard errors in parentheses: heteroscedasticity consistent and adjusted for municipality-level clustering; * : p=0.1, **: p=0.05, ***: p=0.01

Table 11: Simulation Results: 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.

Table 12: Simulation Results: Extended Model

	$\sigma = 3$	$\sigma = 4$	$\sigma = 5$	$\sigma = 6$
	0.08	0.12	0.16	0.18
$(1-\mu) = 0.20$	(25)	(29)	(30)	(31)
	[19]	[22]	[23]	[25]
	0.06	0.09	0.10	0.14
$(1-\mu) = 0.25$	(24)	(28)	(31)	(32)
	[14]	[16]	[17]	[18]
	0.05	0.07	0.09	0.10
$(1-\mu) = 0.30$	(20)	(25)	(27)	(29)
	[10]	[12]	[12]	[12]

Note: Reported numbers are the values assigned to χ, the measure of locational preference heterogeneity, such that $\rho = 3$. Numbers in parentheses are implied percentage real-wage differentials between regions within country A. Numbers in brackets are implied shares of country A's population that prefers not to migrate at the prevailing real-wage differential.