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Protective or Counter-Productive? Labor Market Institutions and the Effect of Immigration on EU Natives

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ABSTRACT

Protective or Counter-Productive? Labor Market Institutions and the Effect of Immigration on EU Natives*

We estimate the effect of immigrant flows on native employment in Western Europe, and then ask whether the employment consequences of immigration vary with institutions that affect labor market flexibility. Reduced flexibility may protect natives from immigrant competition in the near term, but our theoretical framework suggests that reduced flexibility is likely to increase the negative impact of immigration on equilibrium employment. In models without interactions, OLS estimates for a panel of European countries in the 1980s and 1990s show small, mostly negative immigration effects. To reduce bias from the possible endogeneity of immigration flows, we use the fact that many immigrants arriving after 1991 were refugees from the Balkan wars. An IV strategy based on variation in the number of immigrants from former Yugoslavia generates larger though mostly insignificant negative estimates. We then estimate models allowing interactions between the employment response to immigration and institutional characteristics including business entry costs. These results, limited to the sample of native men, generally suggest that reduced flexibility increases the negative impact of immigration. Many of the estimated interaction terms are significant, and imply a significant negative effect on employment in countries with restrictive institutions.

JEL Classification: J23, J61, O52

Keywords: Immigrant absorption, European unemployment, labor market flexibility, entry costs

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Paralleling American interest in the consequences of immigration, recent years have seen increased debate over the impact of immigration in Western Europe. The European immigration debate is fueled by the fact that immigration now accounts for the bulk of population growth in the European Union (EU; OECD, 1999a). Many observers have also noted that increased immigration is likely to be part of any strategy to keep European social security systems solvent. At the same time, the rise in immigration is associated with high levels of anti-foreigner sentiment, and the view that immigrants take jobs from natives is widespread in Europe (Bauer, Lofstrom, and Zimmermann, 2000). The evidence on the employment consequences of immigration in Europe is more fragmentary and harder to assess than the US evidence, which generally shows few effects (see, Friedberg and Hunt, 1995). In a recent survey, however, Bauer and Zimmermann (1999) conclude that, popular sentiment notwithstanding, the employment consequences of immigration for European natives have probably also been modest.[1]

This paper takes a fresh look at the employment consequences of immigration in Western Europe, focusing on a new aspect of the immigration question. A large literature looks at the relationship between labor market institutions and employment levels in Europe. Some observers have argued that persistently high unemployment in Europe is due to institutions that increase turnover and employment costs (see, e g., OECD, 1994). Recently, Blanchard and Wolfers (2000) extended this inquiry by suggesting that the negative employment consequences of a rigid labor market are felt not so much in good times, but rather in the labor market's response to adverse demand shocks. We similarly integrate the debate over institutions with the impact of shocks that occur on the labor-supply side. In particular, we ask whether labor market institutions, such as employment protection, high replacement rates, and business entry costs, affect the employment consequences of immigration-induced increases in the work force.

[1]European country studies include Pischke and Velling (1997) for Germany; Winter-Ebmer and Zweimuller (1997) for Austria, and Dolado, Duce, and Jimeno (1996) for Spain. Similar to Card's (1990) Mariel Boatlift research design, Hunt (1994) compares regions in France after Algerian independence and Carrington and deLima (1996) compare Portugal and Spain after an influx of Portuguese returnees. Hunt's (1992) results show more evidence of disemployment effects than the American studies while Carrington and deLima's (1996) results are inconclusive.

While labor market institutions play a number of roles, one rationale for institutions that reduce flexibility is to protect natives and especially incumbent workers from competition in the labor and product markets. In fact, Rodrik (1997) has argued that the demand for social insurance is in no small measure a response to the pressures of global economic integration, including migration. But the equilibrium consequences of protective regulations and institutions are unclear. We therefore begin with a theoretical discussion of interactions with institutions. Although employment protection and entry barriers may reduce job loss in the short run, our theoretical discussion shows how restrictive institutions can be counter-productive, eventually amplifying the negative employment consequences of immigration for natives.[2]

The main empirical relationship of interest in our study is the effect of immigrant shares on native employment rates, where this effect is allowed to vary with institutional characteristics. Our empirical implementation uses a panel data set for up to 18 European Economic Area (EEA) countries for 1983-99.[3] At our request, the European Commission's statistical agency (Eurostat) produced consistent time series of immigration measures and labor market variables by age, sex, education, and country. These statistics are derived from a series of country-specific labor-force surveys (LFS) using a similar format across countries and over time. This data set allows us to conduct analyses similar to previous immigration research for the US and individual European countries using micro data, while allowing consistent cross-country comparisons. Since European countries define immigrants in terms of nationality, while North American definitions are based on nativity (country of birth), we also briefly compare alternative definitions of immigrant status in Europe, distinguishing non-natives from non-nationals.

In addition to exploring interactions with institutions, our cross-country analysis addresses some of

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[2]Earlier theoretical studies of interactions between immigration effects and labor market flexibility include Schmidt, Stilz, and Zimmermann (1994) and Razin and Sadka (1996).

[3]The modern EU includes 15 countries: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, and the United Kingdom. The EEA includes these plus 3 of 4 states in the European Free Trade Area: Iceland, Liechtenstein, and Norway. Our sample omits Liechtenstein but includes Switzerland, which opted out of the EEA.

the methodological questions raised in the previous immigration literature. First, cross-country data may be less affected by endogenous mobility than within-US data. Second, and most importantly, we experiment with an instrumental variables (IV) strategy, as well as reporting OLS and fixed-effects estimates. The IV estimates use the two 1990s Balkan Wars (in Bosnia and Kosovo) as a source of exogenous variation providing a shock to immigrant flows in Europe. During this period, many immigrants and refugees came from former Yugoslavia. As a consequence, the distance between European population centers and the Yugoslav republics is highly correlated with the wartime proportion of the labor force from non-EU countries, while essentially uncorrelated with the foreign share from EU countries. We implement an estimation strategy based on this variation by using the distance from Sarajevo and the distance from Pristina, interacted with dummies for the war years, as instruments for the immigrant share.

I. Theoretical Framework

We use a competitive model with two types of labor and exogenous separations to illustrate the standard theoretical predictions regarding the impact of that immigrants on natives, and to suggest how the effect of immigrants on native employment might be modified by differences in labor market institutions, such as firing costs. Our theoretical setup is similar to that used in earlier analyses of immigration questions, augmented with elements used by Acemoglu and Angrist (2001) and Saint Paul (1996) to study the effects of labor market regulation and employment protection.

Firm output is assumed to be produced by immigrants and natives with production function,

$$f(\theta_t g_t[N_t, I_t]);$$

$$g_t[N_t, I_t] = (N_t^{\rho} + \gamma I_t^{\rho})^{1/\rho},$$

where N_t is the number of natives (or nationals) and I_t is the number of immigrants (or non-nationals). The variable g_t is a CES-type labor aggregate as in Card (2001); output depends on other factors of production, but this is ignored in the notation. The variable θ_t is an exogenous shifter as in Lalonde and Topel (1991).

The first derivative of the production function, f(·), is positive and the second derivative is negative. Our approach differs modestly from others in the literature in that we have CES interaction between immigrants and natives as a group. In Altonji and Card (1991), for example, immigrant-native complementarity is generated by differences in the skill or education mix of the two groups, with immigrants and natives at the same skill level being perfect substitutes. This approach is harder to motivate in the European context, since immigrants to the EU are not dramatically less educated than natives. Immigrants with the same measured skills may also complement natives because of language or craft skills.

An important feature of many European labor markets is high firing costs. These come in the form of bureaucratic limitations on dismissals, requirements for severance pay, and restrictive collective bargaining agreements. On the other hand, immigrants are probably less likely than natives to be covered by these provisions since immigrants are more likely to work in non-union jobs, on fixed-term contracts (e.g., if they have only temporary work permits), or illegally. We therefore introduce positive firing costs in the amount C_N for natives, with no firing costs for immigrants. Firing costs are paid when, every period, a proportion λ of the labor force becomes unproductive in the current job, and is therefore laid off.[4] Immigrants and natives are paid different real wages, w_{Nt} and w_{It}, with the price of output as numeraire.

Our interest is in immigration effects and interactions with institutions in a simple dynamic setup. We assume price-taking firms act to maximize the present value of profits, with discount factor φ. In this case, firms' objective functions can be written,

$$\Pi = \sum_{t=0}^{\infty} \varphi^t[f(\theta_t g_t[N_t, I_t]) - w_{Nt}N_t - w_{It}I_t - \lambda C_N N_{t-1}].$$

Since adjustment costs are linear, time subscripts can be dropped and the objective simplified:

[4] As in Angrist and Acemoglu (2001), productivity shocks are high enough and firing costs low enough that laying off unproductive workers is indeed worthwhile. In practice, productivity may be endogenous and determined in part by employment protection (see Ichino and Riphahn, 2000, for evidence). We ignore hiring costs since adjustment costs are linear in our model (as in Saint-Paul, 1996).

$$\Pi = \quad (1-\varphi)^{-1}[f(\theta g) - w_N N - w_I I - \varphi\lambda C_N N]. \tag{1}$$

Employment levels are chosen to satisfy the first order conditions

$$f'(\theta g)\theta g_N = w_N + \varphi\lambda C_N = w_N(1 + \varphi\lambda c_N) \tag{2a}$$

and

$$f'(\theta g)\theta g_I = w_I, \tag{2b}$$

where g_N and g_I are derivatives of g[N, I]. Equation (2a), which implicitly defines the demand function for native labor, equates the flow cost of native workers with their marginal product. The flow cost of employing immigrants is just the immigrant wage. Note that firing costs are now defined as proportional to the native wage: $c_N \equiv C_N/w_N$, a ratio that we take to be fixed.

The labor supply of immigrants is assumed to be perfectly inelastic, at least for the range of wage levels observed in the host country. The immigrant population is denoted by M, employed in equal numbers at each of m identical firms, so that $I=M/m$. In contrast, native labor supply is elastic and determined by a second institutional characteristic, unemployment insurance replacement rates, as well as by wages. The distinction between immigrant and native labor supply seems reasonable since natives are more likely than immigrants to have access to social insurance. The native labor supply function is

$$N^* = mN = [w_N(1-r)]^\varepsilon P, \tag{3}$$

where P is the native population, r is the UI or social insurance replacement rate, and ε is the native labor supply elasticity, assumed to be positive. For what follows, it is useful to define the inverse labor supply function, $w_N(N, \varepsilon) \equiv (mN/P)^{1/\varepsilon}(1-r)^{-1}$.

The short-run analysis of this model takes the number of firms, m, to be fixed, while the long-run response is determined by allowing m to be endogenous and determined by the condition that profits are greater than entry costs. In the short run, equations (2a), (2b), and (3) determine the two endogenous wage levels, and the number of employed natives. Since immigrant labor supply is exogenous, the key equilibrium condition can be written as follows

$$\mathit{ln}f'(\theta g[N, I]) + \mathit{ln}\theta + \mathit{ln}g_N[N, I] = \mathit{ln}w_N + \mathit{ln}(1 + \varphi\lambda c_N)$$

$$= (1/\varepsilon)\mathit{ln}(N/P) + (1/\varepsilon)\mathit{ln}m + \varphi\lambda c_N + r. \tag{4}$$

This equation determines native employment, which can then be substituted back into the labor supply equation to find native wages.

Equation (4) provides a basis for the empirical work in this paper. Following Lalonde and Topel (1991), we think of the estimates, which relate the log of native employment to the log of the immigrant share in the labor force, as approximating $\mathit{ln}f'(\theta g) + \mathit{ln}g_N$ to first order in logs.[5] The estimating equation is assumed to hold at the country level since we imagine all firms are identical except possibly for the shift variable, θ, which is assumed to be absorbed by country and year effects. Before turning to the empirical results, we use this theoretical framework to suggest the likely nature of interactions between labor market institutions and immigration. In particular, we are interested in the short-run impact of immigration on total native employment, $N^* \equiv mN$, and the question of whether $\partial N^*/\partial M$ changes with firing costs, c_N; replacement rates, r; and the degree of native wage flexibility. We also consider long-run impacts and the interaction between immigration and barriers to entry.

The short-run employment impact of immigration can be written in elasticity terms as follows,

$$d\mathit{ln}N^*/d\mathit{ln}M = (\partial N/\partial I)(I/N) = \xi_{NI}[\varepsilon^{-1} - \xi_{NN}]^{-1} \equiv e(N, \varepsilon), \tag{5}$$

where ξ_{NI} and ξ_{NN} are the elasticities of factor price for native wage rates with respect to native and immigrant employment. That is, $\xi_{NN}=(\partial w_N/\partial N)(N/w_N)$ and $\xi_{NI}=(\partial w_N/\partial I)(I/w_N)$ along the demand curve for native labor. We use the $e(N, \varepsilon)$ shorthand for $d\mathit{ln}N^*/d\mathit{ln}M$ to highlight the fact that parameters other than the labor supply elasticity modify the immigrant impact through native employment levels. This expression, derived in the appendix, is similar to the corresponding relationship in Johnson's (1980) static model, though in our case, immigration has an ambiguous effect on native employment. While ξ_{NN} is negative, so the denominator is

[5]If there are no firing costs and $f(\cdot)$ is linear, so the production function is constant-returns CES with labor inputs alone, then the left hand side of (4) can be written as a linear combination of $\mathit{ln}N$, $\mathit{ln}I$, and the log of the immigrant share.

positive, ξ_{NI} in the numerator can be positive or negative depending on the extent of immigrant-native complementarity. In a setup like ours, however, immigration is predicted to reduce native wage rates for most plausible parameter values (see, e.g., Altonji and Card, 1991; Ichino, 1993), in which case $e(N, \varepsilon)$ is negative. If immigrants and natives are perfect substitutes (ρ=1), then ξ_{NI} and $e(N, \varepsilon)$ are necessarily negative. Assuming ξ_{NI} is negative, as we do in the discussion which follows, immigration reduces native employment, with larger effects when native labor supply is more elastic.

To see how the employment effects predicted by (5) are affected by changes in firing costs, note that c_N does not appear directly in the derivative $e(N, \varepsilon)$. Thus, any change in employment response is due to the impact of firing costs on employment levels. That is,

$$\partial e/\partial c_N = [\partial e/\partial N][\partial N/\partial c_N].$$

It is clear from (4) that firing costs reduce employment in this model. In particular,

$$\partial N/\partial c_N = -\varphi\lambda/\Delta, < 0,$$

where Δ is (1/N) times the denominator in (5).[6] The scale effect, $\partial e/\partial N$, is likely to be positive, i.e., there is less native job loss due to immigration when N is larger. To see this, it is useful to write out the scale effect as follows:

$$\partial e/\partial N = (\partial\xi_{NI}/\partial N)[\varepsilon^{-1} - \xi_{NN}]^{-1} + \xi_{NI}[\varepsilon^{-1} - \xi_{NN}]^{-2} (\partial\xi_{NN}/\partial N)$$

The wage decline for natives from a given percentage increase in immigrants will likely be smaller (in absolute value) the more natives there are, so the first term on the right hand side is typically positive. The second term is also positive if $\partial\xi_{NN}/\partial N$ is negative. Standard results from demand theory suggest this is a reasonable presumption: in a constant-returns, two-factor model, demand becomes less elastic as factor shares increase, so ξ_{NN} becomes more negative as N grows (see Hammermesh, 1986).

[6]The appendix shows Δ to be $[(1/\varepsilon)N^{-1} - (f''/f')\theta g_N - g_{NN}/g_N] > 0$. Firing costs reduce employment in our model with cross-sectional employment differences, but have an ambiguous effect in partial-equilibrium models with cyclical shocks (as in Bentolila and Bertola, 1990). In a general-equilibrium setting, firing costs also reduce profitability and investment, with consequent job losses (see Hopenhayn and Rogerson, 1993).

These considerations suggest scale effects are indeed positive and therefore the employment loss due to immigration is probably made worse by firing costs. The intuition for this result is straightforward: firing costs reduce native employment levels in our set-up, and reduced employment makes the negative employment consequences of a given number of immigrants worse. Of course, as an empirical matter, scale effects may be small or hard to measure, and, at least in the short run, firing costs are likely to protect incumbent native workers from dismissal.

The analysis of changing *r* parallels the discussion of firing costs since replacement rates do not appear directly in (5). Note that

$$\partial e/\partial r = [\partial e/\partial N][\partial N/\partial r].$$

Similarly, (5) implies that higher replacement rates reduce native employment levels, in this case

$$\partial N/\partial r = -(1-r)^{-1}/\Delta < 0,$$

so higher replacement rates increase any job loss due to immigration. The intuition for this result is the same as for the interaction with firing costs: high replacement rates reduce native employment levels, and reduced employment makes the negative employment consequences of immigration worse.

We omit a detailed analysis of the impact of union wage setting or minimum wages, but look briefly at a stylized model of inflexible wages. Suppose that native wages are fixed at a binding minimum or contract wage, $\bar{w}_N$. Then the effect of immigration on native employment can be shown to be

$$(\partial N^*/\partial M)(M/N^*) = -\xi_{NI}\xi_{NN}^{-1},$$

which is $e(N, \varepsilon)$ with $\varepsilon=\infty$, and is clearly more negative than $e(N, \varepsilon)$ with ε unrestricted. This is the first-order effect; a higher $\bar{w}_N$ also reduces employment, leading to the same sort of scale effect discussed earlier for firing costs and replacement rates. Since the wage-setting mechanism is not specified, our analysis omits any feedback effects whereby union wage demands are moderated as a consequence of competition from immigrants. This possibility is discussed by Schmidt, Stilz, and Zimmermann's (1994).

A. Long-run Effects

Suppose now that the number of firms, *m*, is an endogenous variable eventually determined by the requirement that profits equal entry costs. To see the consequences of endogenous *m* for immigration effects, we first analyze the effect of immigration on profits. The effect of an increase in M on profits with a fixed number of firms is approximately

$$\partial\Pi/\partial M = (1-\varphi)^{-1}\{ -N(\partial w_N/\partial M) - I(\partial w_I/\partial M) \},$$

where N is the equilibrium employment level of natives. This expression is derived by observing that the envelope theorem allows us to ignore terms in $\partial\Pi/\partial N$ and $\partial\Pi/\partial I$. In the short run, increased immigration clearly increases profits in this model because immigrant wages must fall and native wages have been presumed to fall in our previous discussion. As Borjas (1995) notes, the increase in profits due to immigration is generated because, while the last worker hired is still paid value of marginal product, infra-marginal workers are paid less. The exception is if there are constant returns for labor inputs alone, i.e., $f''(\cdot)=0$, in which case profits are always zero and there may be no entry.[7]

Assuming profits were equal to entry costs before immigration and there are diminishing returns to labor inputs, the increase in profits after immigration induces the entry of new firms. Because the entering firms employ additional workers, both immigrant and native, the possibility of endogenous entry reduces and may even eliminate any negative impact of immigration on native employment. To see this, note that with free entry, the effect of immigration on aggregate employment can be shown to be

$$(\partial N^*/\partial M)(M/N^*) = e(N, \varepsilon)[1- \partial lnm/\partial ln M] - \xi_{NN}[\varepsilon^{-1} - \xi_{NN}]^{-1} (\partial lnm/\partial ln M) \qquad (6)$$

where $e(N, \varepsilon)$ is the short-run employment response defined in equation (5). This long-run impact is derived in the appendix. Since $\xi_{NN} < 0$ and $\partial lnm/\partial \ln M \leq 1$, the response with entry is less negative than in the fixed-number-of-firms case and can even be positive. With perfect substitution, i.e., $\rho=1$, the short-run impact of

[7]Even if native wages were to rise due to immigrant-native complementarity, it can be shown that profits increase as long as the overall production function exhibits diminishing returns.

immigration on native employment is necessarily negative, but the long-run impact is zero.

Although entry may eventually raise employment back to pre-immigration levels, in the theoretical medium-run, immigrants will have a diminished, though still negative effect on native employment. So factors that inhibit entry are likely to increase or prolong the displacement of natives by immigrants. Moreover, entry costs probably interact with other wage rigidities, such as firing costs and sticky wages, to aggravate job losses from immigration for two reasons. First, factors that increase labor costs will tend to reduce or slow the entry of firms in response to low-cost immigrant labor. Second, entry costs that reduce native employment levels will interact negatively with other rigidities because of the short-run scale effect.[8] On balance, therefore, our model clearly predicts an association between barriers to entry and native job losses due to immigration.

Overall, the theoretical discussion suggests that the relationship between immigration and native employment is likely to vary across countries according to employment laws, replacement rates, wage-setting institutions, and business entry costs. The impact of unions is hard to measure since most Western Europeans are covered by collective bargaining agreements whether or not they are union members. We therefore focus on interactions with measures of labor market flexibility, replacement rates, and barriers to entrepreneurship. To establish a baseline, however, we begin with a reduced-form analysis of immigration effects that omits interactions with institutional characteristics.

II. Background and Data Description

A. Descriptive Statistics

OECD and other EEA countries have long been host to refugees and economic migrants. In recent years, the European countries with the largest proportion of labor force from non-EU countries have been

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[8]See Bertrand and Kramarz (2001) for recent evidence on the employment consequences of entry costs in the retail industry in France.

Austria, France, Germany, Sweden, Switzerland, and the UK. This can be seen in Table 1, which reports descriptive statistics from the Eurostat labor force surveys for 18 EU and other EEA countries.[9] France and the UK absorb many immigrants from former colonies, while Germany and Austria accept large numbers of migrants from Turkey and Eastern Europe, especially Poland. Sweden has a large foreign population, many of whom come from Middle Eastern countries. Another important supply factor in some countries is the absorption of many ethnically similar migrants. Germany, for example, accepts large numbers of ethnic Germans (known as Aussiedler), mostly from the former Soviet Union. Similarly, roughly 45 percent of the foreign population in Greece in 1997 was of Greek background, mostly from the former Soviet Union and Albania (OECD, 1999a).

Figure 1 documents the time pattern of immigration for many of the important immigrant-receiving countries. The figure plots log counts of foreign inflows relative to the count in base year, 1983. There was little change in immigration through 1988, but the late 1980s and early 1990s saw a marked upturn. The increases were sharpest for Germany, beginning in 1989, with strong increases later in Finland, France, and Sweden. The Benelux countries and Switzerland show a more gradual, hump-shaped pattern. Norway received more immigrants in the 1980s than in the 1990s, while Denmark experienced the largest increase after 1994.

The flow from former Yugoslavia became an important part of the European migration picture after 1990. This flow was generated by the collapse of the Yugoslav state, and especially by the Bosnian and Kosovo wars at the beginning and end of the decade. Figure 2 shows that the number of former Yugoslavian asylum-seekers peaked in 1992, the year that Bosnia-Herzegovina became an independent state and Bosnian Serbs laid siege to Sarajevo, and again in 1999, when NATO launched air strikes on Yugoslavia. Figure 2 also shows that Yugoslavs accounted for more than 30 percent of all asylum-seekers in the war years. Yugoslav asylum seekers were a significant part of the total foreign inflow, with wartime modes at 10-15

[9]Information on data sources and extracts is provided in the data appendix.

percent. Since many foreigners in EU countries come from other EU countries, the effect of the Yugoslav asylum seekers on the non-EU foreign share is considerably larger than indicated by the figure. Our data show that in the 1995-99 period, for EU countries with information on immigrants' country of origin, roughly one-third of male immigrants aged 20-59 from non-EU countries were from former Yugoslavia.

The importance of migration within the EU is also documented in Table 1. In some countries, many non-national residents in the labor force are from other developed European nations. Luxembourg is a clear outlier, with over one-third of its labor force from other EU countries. France is more typical, with 2-3 percent of its labor force from other EU countries. In the analysis below, we distinguish between EU and non-EU foreigners, and use this distinction to control for intra-EU migration that potentially responds to the number of non-EU immigrants.

B. Nationality Versus Nativity

The immigration statistics in Table 1 are based on a distinction between home-country nationals and resident non-nationals, the latter group consisting of what Europeans typically call "foreigners" and what Americans refer to as "resident aliens". Most discussions of immigration in Europe use an immigrant definition based on nationality. Thus, ethnic Germans moving to Germany are not counted in a nationality-based definition of immigrants, while any immigrants who become naturalized citizens cease to be identified as immigrants on a nationality basis. In contrast, in Australia, the US, and Canada, countries with a long tradition of immigrant absorption, immigrants are usually defined by nativity; i.e., an immigrant is any foreign-born resident whether or not a citizen of the host country. Countries defining immigrants by nativity tend to have higher naturalization rates than countries using nationality definitions.

The LFS data allow us to explore the overlap between alternative definitions of immigrant status in

1996.[10] The first two columns of Table 2 report the proportion of non-nationals who are foreign born, separately for EU and non-EU nationals. The statistics in the table are for men and women aged 20-59. These columns show that in all countries for which we have data on country of birth, almost all of those with non-EU nationality were foreign born. Thus, among adults, most non-nationals are indeed immigrants.[11] On the other hand, columns 3 and 4 of Table 2 show that many foreign born residents of EEA countries became naturalized citizens, or were granted citizenship on the basis of ethnicity at the time they arrived in the host country. In France, for example, about half of the residents born in a non-EU country obtained French citizenship. This figure probably includes many people from former French colonies, including people of French ancestry or mixed descent.

Naturalization rates for non-Europeans are surprisingly high overall, but the last two columns show that few *recent arrivals* were naturalized. These columns report naturalization rates for foreign born residents who arrived in the last five years, separately for EU and non-EU countries of birth. For example, for those born in non-EU countries, the recent-arrival naturalization rate is essentially zero in France, 9 percent in Belgium, and 7 percent in the UK. This is not surprising since many EU countries impose an extensive waiting period before naturalization is possible.

In many countries, naturalization rates are lower for recent arrivals born in other EU countries than for those born outside the EU. This probably reflects increasing mobility between EU member states, and the fact that EU citizenship in any country already grants many citizenship privileges in other EU countries. Another interesting feature of the table is the relatively high naturalization rate for recent arrivals in Germany and Greece. As noted earlier, this reflects a preponderance of same-ethnicity migrants coming from the

[10]We have data on country of birth only for 1992-99. Statistics for 1996 are typical of this period. A shortcoming of the LFS data is that the national sampling frames are mostly limited to private households, therefore missing immigrants in relocation camps and hostels. Immigrants working illegally are probably also unlikely to respond to the LFS.

[11]Most EU countries now have provisions for the naturalization of native-born children of non-nationals at age 18. For example, Germany's 1990 Act of Naturalization "specifically extended naturalization to young foreigners who have lived in Germany for a long time and wish to remain there." (OECD, 1995, p. 166).

former Soviet Union and (in the case of Greece) Albania. On balance, however, the statistics in Table 2 suggest that for most countries, the group of non-nationals can be seen as roughly coincident with the group of recently-arrived foreign born residents. We explore the relationship between these two immigrant definitions further when estimating immigration effects.

C. Immigrant Employment: Policies and Descriptive Statistics

Before turning to the estimates, we briefly discuss government policies and a few statistics relating to the ease of labor-market access and legal employment opportunities for European immigrants. The OECD (e.g., 1999a) migration volumes describe aspects of immigration policies in the EU. Of special relevance for our study is the treatment of immigrants and asylum-seekers from Yugoslavia since we use the Balkan Wars as a source of exogenous variation in immigrant flows. Official policy appears to allow many of these people to work, at least around the time of the Bosnia war.

In Austria, which absorbed 100,000 Bosnian refugees between 1992 and 1995, the majority of Bosnians have a long-term work entitlement. Germany has made it more difficult to obtain asylum since 1993, but the largest number of asylum seekers come from former Yugoslavia, and many work permits were apparently issued to asylum seekers and other foreigners, especially in 1994 and 1995. Italy saw a tripling of foreign employment between 1990 and 1997, a large decline in unemployment among foreigners, and a substantial expansion of service-sector employment fueled by immigrants, in spite of the fact that an estimated 89% of asylum seekers entered Italy illegally in recent years. In Sweden in 1998, the largest immigrant group in the labor force after the Fins were those from former Yugoslavia. A special visa program for Bosnians and other parts of the former Yugoslavia operated in Sweden from 1993-96. The largest number of migrants to Switzerland between 1994 and 1997 also came from former Yugoslavia; the largest group of non-EU workers in Switzerland in this period were Yugoslavs.

Participation rates for immigrants are typically below those for natives, but most immigrant men aged

20-59, including those from former Yugoslavia, are in the labor force. This can be seen in Table 3, which reports labor-force participation rates for natives and non-EU immigrants, with statistics for immigrants shown separately for those arriving in the past 5 years (recent arrivals) and those arriving earlier (veteran immigrants). The table reports statistics averaged for all available years from 1995-99 since this is the period when LFS coverage was broadest.

Participation rates for veteran male immigrants are generally close to those for natives, and even higher than for natives in Italy, Austria, Greece, and Spain. As many other researchers have found, our data generally show recent immigrants have lower participation rates than both natives and veteran immigrants. On the other hand, the majority of male immigrants count themselves as in the labor force in every country. The contrast between natives and recent immigrant is greater for women than for men in many countries. It is also worth noting that the participation rates for Yugoslavs are similar to those of other non-EU immigrant groups in most countries. In Sweden, for example, 79 percent of veteran Yugoslav men and 81 percent of other veteran non-EU men were in the labor force, while 65 percent of recent Yugoslav men and 62 percent of other recent non-EU men were in the labor force.

Unemployment is typically higher among immigrants than natives, so the difference in employment rates by nativity generally exceeds the difference in participation rates. This can be seen in panel B of Table 3, which reports employment-to-population ratios. Most veteran immigrant men were working, and employment rates in most countries were similar for veterans who immigrated from former Yugoslavia and those who immigrated from other non-EU countries. The employment rates of recently arrived Yugoslavs were lower than those of veteran Yugoslavs, ranging from 27 percent for men in the UK, to rates above 75 percent in Austria, Greece, Luxembourg, and Spain. Again, the employment pattern of recent arrivals from Yugoslavia are typically similar to those of other recent immigrants from non-EU countries, though there are exceptions. The officially measured employment rates for immigrants generally increase with time in country. And, as we noted earlier, the distinction between officially-measured employment and labor force participation may not be clear-cut for immigrants.

Also relevant are the comparative skill and education levels of immigrants and natives, since there is probably more competition for jobs within groups than between groups. Our LFS extract includes information on the size of three schooling groups, categorized by International Standard Classification of Education (ISCED) levels 0-2, 3-4, and 5 and above. ISCED level 2 typically denotes a ninth grade education, and corresponds to the end of compulsory education in many countries. We therefore define the "low-educated" as those with ISCED levels 0-2 and compare proportions in this group.

Averaging data for 1995 and 1999, about 33 percent of native men in our sample are in the low-education group.[12] The proportion is higher for veteran immigrants from non-EU countries, about 42 percent, with the proportion low-educated similar among recent non-EU immigrants at 40 percent. Recent immigrants from Yugoslavia are similarly less-educated, with about 39 percent in the low-education group. The proportion with low education is even higher among veteran male immigrants from Yugoslavia, roughly 48 percent. Almost half of non-EU female immigrants are low-educated (47 percent); likewise for recent Yugoslav women. The least educated group consists of veteran female immigrants from Yugoslavia, where 65 percent are less-educated. The immigrant/native contrast in schooling levels is larger for women than for men since only 37 percent of native women are in the low-education group. Overall, however, the education statistics show considerable overlap between the immigrant and native schooling distributions.

III. Estimates of Immigration Effects

A. OLS Estimates

The equation of interest links the employment rates of natives with the immigrant proportion in the labor force. We begin with a model that imposes a constant immigration effect across all countries, turning afterwards to models that allow for interactions with institutions. The first equation estimated is

[12]The sample used to compare schooling levels omits Germany, Iceland, and Switzerland.

$$ln(y_{ijt}) = \mu_i + \delta_t + \beta_j + \alpha_i ln(s_{jt}) + \varepsilon_{ijt} \tag{7}$$

for demographic group i, country j, and year t. The model includes country and year effects, β_j and δ_t, with group main effects included when demographic groups are pooled. The regressor $ln(s_{jt})$ is the log of the immigrant share and the dependent variable is the log of the employment-to-population ratio for natives. The immigrant share is defined as either the proportion of the labor force made up of non-nationals from non-EU countries, or the proportion of the labor force born in non-EU countries.[13] As noted above, we think of this equation as approximating the first-order condition determining native employment. This can also be seen as a general reduced-form relationship between native employment and the immigrant share. In either case, the most important omitted variables are time-varying productivity or labor demand shocks correlated with both immigrant shares and native employment.[14]

The first set of estimates, reported in Table 4, use data for 1992-99 only. This table briefly explores the consequences of switching from nationality-based to nativity-based definitions of immigrant status. For example, the first row shows a negative but insignificant relationship between the non-national share in the labor force and male nationals' employment rates. The estimates remain insignificant when pooled across age and schooling groups, when estimated separately for young and old workers, and when estimated in three separate schooling groups. Estimates using nationality definitions are likewise insignificant for women, though less precise as well.

The second panel in Table 4 shows that switching from a nationality-based to a nativity-based definition of immigrant status fails to generate any evidence of an employment-immigration relationship. On the other hand, the third panel in the table, which replaces the non-native share with the proportion of

[13]For the purposes of this distinction, we define "EU countries" on a time-invariant basis though EU membership was changing. See the appendix for details.

[14]We look at the employment effects of immigrants in the labor force, i.e., working or unemployed, as opposed to measuring immigrants as a proportion of population. This seems like a conservative approach to measuring the size of the immigrant group that can affect natives, since immigrants who can not legally work may nevertheless work illegally, while still reporting themselves as out of the labor force in surveys. Unemployed immigrants may also put downward pressure on native wages.

the labor force consisting of immigrants who arrived in the last five years, shows a negative and significant effect on the employment rates of less-educated men. This finding is consistent with the view that recent arrivals are more likely than veteran immigrants to undercut less-skilled natives in the labor market. Veteran immigrants probably behave more like natives, with higher wages, and a higher probability of coverage by labor laws and collective-bargaining agreements.[15] The magnitude of the effects for less-educated men is such that an increase in the recent immigrant share of 10 percent would reduce native employment rates by less than a third of a percent.

The remainder of the analysis uses a nationality-based definition of immigrant status because nationality variables are available in our sample beginning in 1983, while the country-of-birth variables start in 1992. The longer period generates more precise estimates, facilitates control for country trends, and allows us to implement an instrumental variables strategy based on the Balkan wars. It should also be noted that the Balkan war instruments may disproportionately capture the effects of recent arrivals. Of course, recent arrivals have lower labor force participation than veterans, but our endogenous regressor includes only those immigrants who are in the labor force.

Table 5 reports OLS estimates for the 1983-99 sample period, with additional estimates for 1992-99 for comparison. In addition to including more data, the model used to construct the estimates reported in Table 5 differs from equation (7) in that it also includes the share of the labor force from EU countries. That is, in addition to the log of the non-EU foreign share, $ln(s_{jt})$, the estimating equation also includes the log of the foreign share with EU nationality outside the country of residence, $ln(u_{jt})$. As noted earlier, many EU countries have a large number of foreign residents from other EU countries. These internal EU migrants are probably similar in some respects to US internal migrants, in that they are drawn to host countries by job assignments and employment opportunities.

[15]Card (2001) similarly argues that recent immigrants may be more likely than veteran immigrants to compete with less-skilled natives.

Figure 3 plots one non-national share variable against the other, after removing country and year effects. This figure shows that the EU and non-EU foreign shares are, in fact, positively correlated. This finding is similar to Pischke and Velling's (1997) results for Germany, which show that immigrants and natives seem to be attracted to the same locations. In our case, the positive correlation in the two foreign share variables suggests that migration within the EU does not act to "undo" the effects of non-EU migrants. We nevertheless include the EU share variable in some of the estimating equations as a proxy to control for omitted variables that might bias the OLS estimates. These omitted variables include host-country economic conditions that are likely to attract migrants.

Not surprisingly, the results for 1992-99 in Table 5 are similar to those for the same period shown in Table 4, again offering no significant evidence of an immigrant-employment relationship. For the 1992-99 sample, the only change in specification from Table 4 is the inclusion of the EU share variable, which is positively correlated with native employment. This positive correlation, however, is unlikely to be causal, and probably reflects the "pull" exerted by employment opportunities in one EU country on other EU residents.

Extending the sample back to 1983 changes the story somewhat. First, note that we no longer report estimates by schooling group in the extended sample because the schooling variables are unavailable in the LFS in the earlier period. Second, pooling the young and old age groups leads to a negative and statistically significant effect on male native employment of -.021. The difference between this result and that for the 92-99 sample is partly because the effect is bigger and partly because it is estimated more precisely. Estimates using young men show an even larger negative effect of -.037. Taking the pooled significant estimate as representative of the impact on men, the magnitudes are such that 100 immigrants in the labor force cost about 35 native jobs in a country where 5 percent of the labor force is non-EU foreign (as Table

1 indicates for Germany in 1989).[16] In contrast, the estimates for younger women using 1983-99 data show no significant effects.

Equation (7) relies on time-invariant country effects to control for omitted variables correlated with immigration rates. OLS estimates of the parameters in this equation are biased if immigration is correlated with omitted time-varying variables or country-specific trends, a problem that is likely to be worse with a longer time-series sample. We therefore also report results replacing the country effect, β_j, with a country-specific linear trend, $\beta_{0j} + \beta_{1j}t$. The negative employment effects for younger women become slightly larger and statistically significant in models with country trends, but the results for younger men, while still negative, are smaller, less precise, and no longer significantly different from zero. The estimates pooling age groups are negative but insignificant for both men and women in models with country trends.

B. IV Estimates

The OLS estimates for 1983-99 are all negative, though not significant in most of the models that control for country trends. But the OLS estimates may be biased upwards by immigrants choosing to locate in countries where their employment prospects are best. This section discusses estimates of immigration effects in the 1983-99 sample using an IV strategy. The choice of instruments is motivated by Figure 2, which shows a sharp run-up in the number of Yugoslavs among European immigrants in the early and late 1990s. This figure suggests that distance from the Yugoslav conflict should be a good predictor of the foreign share in the 1990s.[17]

The first-stage equation for the IV estimates is

$$ln(s_{jt}) = \tau_t + \psi_j + b_{jt}\pi_b + n_{jt}\pi_n + k_{jt}\pi_k + \eta_{ijt} \tag{8}$$

[16]We use the German foreign share as a base even though this exceeds the median foreign share in our data since the LFS misses some types of foreign workers. As noted above, the LFS sampling frames are limited to private households and the LFS probably also misses many illegal, temporary, seasonal, and cross-border workers. See OECD (1995) for a discussion of the foreign worker undercount.

[17]Schwartz (1973) is an early study documenting the impact of distance on migration.

where

b_{jt} = distance from Sarajevo × dummy for 1991-95 (Bosnia War years)

n_{jt} = distance from Sarajevo × dummy for 1996-97 (Inter-war years)

k_{jt} = distance from Pristina × dummy for 1998-99 (Kosovo War years)

are the excluded instruments, and τ_t and ψ_j are period and country effects. The distance from potential host countries is measured either from the nearest city with a population of at least 100,000 or from the capital. In some specifications, we add the log of the EU foreign share (i.e., $ln(u_{jt})$) to this equation as an exogenous covariate. In practice, as we show below, this has little effect on the IV estimates because the EU share is largely uncorrelated with the instruments.

The essence of the IV strategy is to look for a break in the time-series behavior of employment rates for countries relatively close to Yugoslavia. Therefore, as a specification check, we also estimated models with a parametric controls for linear country-specific trends, as in the OLS models with trends in Table 5. The specifications with country trends replaces ψ_j with $\psi_{0j} + \psi_{1j}t$ in the first stage, in which case a corresponding term is also included in the second stage ($\beta_{0j} + \beta_{1j}t$) as an additional exogenous covariate.

Conditional on country and year effects, distance from the former Yugoslav republics is associated with a sharply lower immigrant share in the war years. This can be seen in panel A of Table 6, which reports the coefficients on b_{jt}, n_{jt}, and k_{jt}, plus a pre-war interaction as a specification check. The coefficients are scaled so that they represent the effect of 1000 miles. Thus, the differential distance from Graz, Austria to Liege, Belgium, about 500 miles, reduces the non-EU foreign share during the Bosnia War by 30-40 percent (see columns 1-4 in the table).

The pattern of estimates in models without trends is consistent with the notion that immigration was highest during the war years, with a moderate decline in the inter-war years. Adding country trends changes the pattern somewhat, but the estimates are not precise enough for the change to be statistically meaningful. It is perhaps to be expected that the inter-war reduction is not sharp since the inter-war and Kosovo war dummy are also correlated with the presence of Yugoslavs who stayed in their host countries. Importantly, however, the estimates in the last row in panel A shows that adding a dummy for pre-war years to the set of interactions generates no evidence of a pre-existing immigration trend associated with distance from Sarajevo for either distance measure.

As an additional check on the first stage, panel B of Table 6 shows the results of replacing the non-EU foreign share with the EU foreign share as the dependent variable in equation (8). That is, we replace $ln(s_{jt})$ with $ln(u_{jt})$. These estimates show no relationship between wartime interactions with distance to Sarajevo or Pristina and the EU share. This is encouraging since it suggests the estimates in panel A are really picking up the effect of immigrants from former Yugoslavia. Moreover, it means that when implementing the IV estimation strategy we can reasonably choose to ignore the possibly endogenous covariate, $ln(u_{jt})$, since this is essentially uncorrelated with the instruments. Finally, note that the first-stage estimates are generally similar whether distance is measured from capital cities or big cities. But because the first-stage relationship is stronger when distance is measured from large cities, we used this variable to construct the second-stage estimates discussed below.

The 2SLS estimates using b_{jt}, n_{jt}, and k_{jt} as instruments are reported in Table 7, separately for models that do and do not control for $ln(u_{jt})$. For men, the estimated effects are on the order of -.05 when age groups are pooled and -.08 for those under 40. IV estimates for men in models that include country trends are smaller, though still significant for those over 40. Adding the EU share as a control variable has little effect on the estimates in models that do not include country trends. In models with country trends, the estimates including the EU share are larger, but this turns out to be due to the fact that the estimates with country trends are sensitive to the change in sample (from 420 to 402 observations) when the EU share is included.

The IV estimates for men are consistently negative and at the upper end of elasticity estimates reported by Borjas (1994). As in Card (2001), the IV estimates are larger in magnitude than the corresponding OLS estimates. Taking -.05 as representative, the IV estimates for men predict that increasing the foreign share by 10 percent would reduce employment by half of a percent in a country where 5 percent of the labor force is foreign. On a per-worker basis, this implies that 100 immigrants in the labor force cost about 83 native jobs, a large effect in levels. Such significant displacement, especially when compared with results from natural experiments such as the Mariel Boatlift, could be due in part to the interactions described in the theoretical section. It may be that immigrants have a greater displacement effect in Europe than in America because of differing institutions. It should also be noted, however, that the estimates with country trends are mostly smaller and not significant.

While many of the 2SLS estimates for men are imprecise, they suggest a pattern of negative and reasonably stable effects. The results for women are harder to interpret. On one hand, the 2SLS estimates for women show very large negative effects, clearly too large to be attributable to the effects of immigrants. On the other hand, the estimates are greatly reduced by controlling for country trends. The coefficient on the exogenous EU share also falls sharply when trends are added to the models for women. In models with trends, the estimated effect of the non-EU share on the employment of young women is similar to that for men, while the estimates for older women are still larger. Note that percentage effects for women translate into effects on levels that are about half the size of those for men because of lower female labor force participation rates.

The marked sensitivity of the estimates for women to the inclusion of country trends suggests these estimates are probably driven by forces other than increased immigration. One problem with the IV strategy for women is that some countries saw dramatic changes in female labor force participation (LFP) over this period while female LFP in other countries was more stable. In Italy, for example, employment to population ratios of prime-age women increased by only 3.3 percentage points between 1990 and 1998. Similarly, in Greece, female employment rates rose by just 4.5 percentage points. In Belgium, in contrast, which is much further away from Yugoslavia, female employment rates rose by 8.4 points. Similarly, in distant Ireland, female employment rates rose by 17.5 points. This sort of contrast in female employment growth probably induces a spuriously large IV estimate of immigration effects, since the trend growth was typically larger in countries farther from Sarajevo and Pristina. For men, on the other hand, employment rates have been more stable, with less evidence of trends that differ sharply by country or region.[18]

IV. Immigrants Interact with Institutions

In this section, we turn to an empirical investigation of the main question raised in the theoretical discussion in Section II: how does the impact of immigration on native employment vary with the restrictiveness of labor market institutions? In particular, do institutions that make labor markets more rigid or less competitive change the employment consequences of immigration for natives? The theoretical section suggests that restrictive labor standards that affect natives more than immigrants are likely to aggravate any job losses from immigration, though firing costs may protect incumbent native workers from dismissal, at least in the short run. Higher replacement rates improve natives' non-work options, reducing employment levels, and therefore increasing native job loss. Reduced wage flexibility worsens the employment impact of immigrants because of scale effects and because rigid wages make native workers less competitive with immigrants. Finally, higher entry costs are also predicted to amplify the negative effects of immigrants on natives since new firms create jobs that would otherwise tend to neutralize any displacement effects. And entry costs that reduce employment levels also have a scale effect.

The OLS estimates of immigration effects discussed in the previous section may be biased towards zero because of endogenous migration, while some of the IV estimates are probably too large to be due solely to immigration, especially for women. This may be a consequence of omitted trends correlated with the instruments. We therefore continue to present results that control for country trends, as well as OLS estimates, which arguably provide a lower bound even if the instruments are contaminated. Our empirical strategy looks at OLS and IV estimates of interactions with measures of three of the institutional features

[18]The statistics quoted in this paragraph are for women aged 25-54 from Table C in OECD (2000). Employment trends in our data are similar.

discussed in the theory section. The first is a summary index of labor standards that reflects the extent of employment protection, restrictions on work hours and employment contracts, administrative or union oversight in hiring and firing decisions, and minimum wages. The second is the average replacement rate. Both measures were taken from Table 4 in Nickell (1997), and are repeated here in the data appendix. Labor standards are captured by an index ranging from 0-7, with 7 denoting the most restrictive institutions. Replacement rates are measured in percent, ranging from 20-90 in our sample. Finally, we explore interactions with a measure of entry costs taken from Nicoletti, Scarpetta, and Boylaud (2000). This is an index of barriers to entrepreneurship ranging about .5- 2.75 in our sample (reproduced in the appendix).

A. Estimates of Interaction Effects

The equation used to estimate interactions between immigrants and labor market institutions is

$$ln(y_{ijt}) = \mu_i + \delta_t + \beta_j + [\alpha_{0i} + \alpha_{1i}\tilde{x}_j]ln(s_{jt}) + \nu_{ijt}, \qquad (9)$$

where $\tilde{x}_j$ is a variable characterizing institutions, measured as the deviation from the median institution value among the countries for which we have data on institutional characteristics. The parameter α_{0i} therefore captures the effect of immigration on demographic group i in countries with the median institution value, while the interaction term, α_{1i}, describes how this effect changes as institutions change. We therefore think of α_{1i} as the derivative of $e(N, \varepsilon)$ with respect to institutional variables. When $ln(s_{jt})$ is treated as endogenous, the instrument list used to estimate equation (9) is augmented with interactions between $\tilde{x}_j$ and the instruments used to estimate equation (7). Note that this setup fails to identify the effects of the institutions themselves since our institution variables are time-invariant and absorbed by country effects.[19]

The analysis of institutions is limited to the sample of men since the 2SLS estimates for women are considerably more sensitive to control for country trends. As noted earlier, we believe this sensitivity is due

[19]A recent OECD (1999b) study of employment protection shows a negative correlation between measures of protection and the employment of youth and prime-age women, with a positive relationship for prime-age men.

to the strong country trends in female labor force participation. These trends appear to vary across countries in a manner correlated with distance from the Yugoslav conflicts.

The estimates of equation (9) for men suggest that immigration effects are probably more negative in countries with less flexible labor markets, higher replacement rates, and higher entry costs. The estimates, reported in Table 8 for specifications without country trends, begin with an analysis of institution interactions one at a time. The table reports estimates of α_{0i}, labeled "Main Effect" and α_{1i}, labeled "Interaction". The first column shows OLS results for men in both age groups. The interaction with labor standards in this specification is estimated to be a statistically significant -.0077, indicating that increasing the index of labor standards' severity by one unit would increase the negative effect of immigration from -.027 at the median to about -.035. The interaction terms are larger for young men than for men over 40. Similarly, the pooled interaction with replacement rates is -.0009, so a 10 percentage point increase in replacement rates would increase the negative effect of immigration from -.027 at the median to -.036. The results of including both labor market interactions are reported in Panel B of the table. Including both interactions generates OLS estimates that are similar to, though somewhat smaller than, the estimates generated by including the interactions one at a time. Again, the effects are larger for older men.

The 2SLS estimates for men, reported in columns 4-7, differ from the OLS estimates in that both the main effects and interaction terms are less precisely estimated. The OLS and 2SLS estimates of main effects are similar for models that include interaction terms with labor standards. The 2SLS estimates of the interaction terms are much larger than the corresponding OLS estimates, however, and again negative and significant. The 2SLS estimates of interaction terms with replacement rates are not significant, and the 2SLS estimates of main effects in models with replacement rate interactions are not significant. On the other hand, the 2SLS estimates of models incorporating interactions with both labor standards and replacement rates show significant negative interaction terms for labor standards in the pooled and young-men specifications, with insignificant negative interactions for replacement rates in all specifications.

As with the measures of labor market flexibility, the results of estimating models allowing interactions with entry costs show immigration effects that are more negative in countries with barriers to entry. For example, the pooled estimate in column (1) suggests that increasing entry barriers by .5, the difference between Sweden and Denmark, would increase the negative effect of immigration from -.027 at the median to -.043. The interactions with entry costs are again larger for men under 40, and larger when estimated by 2SLS. The results of our effort to estimate the interaction with entry costs jointly with interactions with labor market flexibility are less clear cut. OLS estimates of interactions with entry costs are similar in magnitude to the results when estimated one at a time, though no longer significant. Interactions with replacement rates also remain significant and negative in this specification. The corresponding 2SLS estimates of interactions with entry barriers are much larger (i.e., more negative), perhaps implausibly so, though the interaction-term standard errors are also large.

While our ability to distinguish specific institutional mechanisms is limited, the results show a pattern of negative interactions between immigration and institutional variables that reflect reduced labor and product market flexibility. As a further check on the results with institution interactions, we added country trends, as in the non-interacted models discussed earlier. This necessarily leads to a loss of precision since the interactions with institutions are correlated with country-specific linear trends. The results with country trends, reported in Table 9, nevertheless show some evidence of negative interactions, and generate no significant positive estimates. For example, the OLS estimates of interactions with replacement rates are negative and significant, as is the 2SLS estimate of the interaction with replacement rates for older men. These results hold up when replacement rates and labor standards are entered jointly and with entry barriers. The 2SLS estimates of interactions with barriers to entrepreneurship are also negative and significant when this is the only interaction term.

V. Summary and Conclusions

This paper presents new evidence on the question of how immigration affects native employment. The empirical analysis uses a cross-country panel for European countries with data on employment and immigration by demographic group since 1983. We are especially interested in exploring the extent to which any displacement effects of immigration are mitigated or amplified by reductions in labor market flexibility and business entry costs. Our analysis of immigration effects in Europe is facilitated by the availability of comparable employment and immigration statistics for different countries and demographic groups. We also exploit instrumental variables derived from the immigration shocks caused by the two Balkan wars.

The estimates reported here typically show that an increase in the foreign share of 10 percent would reduce native employment rates by .2 to .7 of a percentage point. The OLS estimates are at the low end of this scale, while the IV estimates are mostly larger than the corresponding OLS estimates, implying substantial displacement of native workers by immigrants. Such large effects could be explained by the institutional mechanisms outlined in our theoretical discussion and by the fact that there has been remarkably little employment creation in most of Western Europe in the last two decades, while immigrant employment has grown considerably. Since many immigrants work, their jobs may well have come at the expense of natives. Of course, effects this size may also signal problems with the IV identification strategy, and the IV estimates are not very precise. It therefore seems reasonable to interpret the OLS and IV estimates reported here as bracketing the true effect.

Our main focus is on the question of whether measures of labor and product-market flexibility change the impact of immigration on native employment. This question has important policy ramifications since many Western European countries are preparing to accept more immigrants, and working to integrate previous immigrant cohorts more fully. Though restrictive institutions can play a protective role, our theoretical framework suggests that institutions such as firing costs, high replacement rates, rigid wages, and business entry costs, will likely aggravate the negative impact of immigration on equilibrium native employment. Part of this interaction is due to scale effects: institutions that reduce employment levels will tend to make the effect of a given number of immigrants worse. Higher entry barriers and reduced wage flexibility also have a direct effect that increases the impact of immigrants on native employment.

The empirical results offer some support for the view that reduced flexibility makes immigrant absorption more painful, at least when viewed from the perspective of native employment. Models that allow the impact of the foreign share on the employment of native men to vary with an index of labor market flexibility, replacement rates, and entry costs tend to show larger immigration effects when flexibility is reduced and replacement rates and entry costs increased. These negative interactions are apparent in the OLS and many of the IV estimates, though the IV estimates of interaction terms are less precise, especially when country trends or more than one institutional characteristic is included in the model. While specific channels are difficult to identify, the results suggest that reduced labor market flexibility fails to protect natives from job loss due to immigration, and may make immigration-related job losses worse.

APPENDIX

1. Data

Data for Figures 1 and 2

The data plotted in Figure 1 are from OECD (1999) and earlier volumes in the same series. The data plotted in Figure 2 are from http://www.unhcr.ch/statist/99oview/toc.htm (Refugees and Others of Concern to UNHCR - 1999 Statistical Overview, published by the United Nations High Commissioner for Refugees, Geneva, July 2000).

The Eurostat Labor Force Survey

The Eurostat LFS data set is documented in Eurostat (1998) and in a variety of memos released with these data. The LFS surveys are carried out by national statistical agencies according to guidelines issued by the European Community. The sampling frame in all countries covers only private households and not group quarters. This is probably important for the coverage of immigrants in some countries. Sampling rates, sample sizes, and interview methods (e.g., use of CATI/CAPI) vary from country to country. The LFS samples are stratified in a variety of ways, but the sample statistics we received from Eurostat were already weighted to population counts. We used these population weights to aggregate cell statistics where necessary (e.g., to combine age groups). Our estimates treat country statistics as population parameters, that is, we did not weight to adjust for differences in country size. We experimented with alternate weighting schemes and found weighted-by-population estimates to be similar. Response rates vary from a low of 55-60 percent in the Netherlands to 98 percent in Germany, with the median response rate at 87 percent. Labor force status is defined using a consistent definition based on “actual status in the reference week”. We checked data quality and our processing by comparing statistics we constructed with those published in the OECD (2000) *Employment Outlook*.

Time-consistent Definition of immigrants' EU status

The analysis here distinguishes non-natives and non-nationals according to home country membership in the EU. This distinction is complicated in practice by the fact that the list of EU member countries changed a number of times over the sample period, and home countries are not always identified separately. Moreover, we cannot distinguish East from West Germany in the pre-unification period. Our analysis uses a time-consistent definition of EU membership, though this definition differs slightly for the 1983-99 and 1992-99 sample periods. For 1983-99, we define the EU as the original EU-12 plus Austria and Norway. All immigrants from Germany are defined as EU (the results are not sensitive to the classification of German immigrants as being from a non-EU country). For 1992-99, we define the EU as the EU-15 plus Norway, Iceland, and Switzerland.

Cell Statistics and Confidentiality Edit

At our request, Eurostat provided us with tabulations of LFS cell statistics by country, year, age, schooling group, nationality, and nativity. Research data provided by Eurostat are released with the stipulation that cell statistics below country-specific thresholds not be released or used in statistical analyses. The results reported here use only those cells above the disclosure thresholds, as determined by a table provided by Eurostat. The restrictions we used are those from "column A" of the Eurostat Guidelines table, as described in the latest release of the "New Chronos" data set.

Distance and Labor-Market Flexibility variables

Distance measures are as-the-crow-flies distance in miles from Sarajevo and Pristina to the capital city or the nearest city with a population over 100,000. Replacement rates and the labor standards index are from Nickell (1997; Table 4). Barriers to entrepreneurship are from Nicoletti, Scarpetta, and Boylaud (2000; A3.2).

Country	Distance from				Replacement rate	Labor standards	Entry barriers
	Sarajevo		Pristina				
	Big city	Capitals	Big city	Capitals			
AT	262.5	318.8	412.5	450.0	50	5	1.60
BE	759.4	806.3	909.4	965.6	60	4	2.55
CH	534.4	571.9	693.8	731.3	70	3	2.24
DE	440.6	712.5	590.6	862.5	63	6	2.10
DK	862.5	853.1	993.8	965.6	90	2	1.32
ES	843.8	1162.5	975.0	1293.8	70	7	1.77
FI	1143.8	1143.8	1209.4	1209.4	63	5	1.93
FR	553.1	834.4	703.1	993.8	57	6	2.73
GR	318.8	487.5	168.9	356.3			1.66
IE	1293.8	1293.8	1443.8	1443.8	37	4	1.20
IS	1903.1	2081.3	2043.8	2231.3			
IT	234.4	328.1	243.8	450.0	20	7	2.74
LU	693.8	693.8	862.5	862.5			
NL	834.4	853.1	984.4	1003.1	70	5	1.41
NO	1153.1	1153.1	1256.3	1256.3	65	5	1.33
PT	1387.5	1481.3	1537.5	1612.5	65	4	1.46
SE	843.8	1059.4	956.3	1153.1	80	7	1.80
UK	1003.1	1003.1	1162.5	1162.5	38	0	0.48

2. Derivations

Derivation of short-run employment change, (5)

Differentiate the equilibrium condition, (4), with respect to I:

$$[f''/f']\theta[g_N(\partial N/\partial I) + g_I] + [g_{NN}^*(\partial N/\partial I) + g_{NI}^*)] = (1/\varepsilon)N^{-1}(\partial N/\partial I)$$

where $g_{NN}^* = g_{NN}/g_N$ and $g_{NI}^*=g_{NI}/g_N$. Therefore

$$\partial N/\partial I = [(f''/f')\theta g_I + g_{NI}^*][(1/\varepsilon)N^{-1} - (f''/f')\theta g_N - g_{NN}^*]^{-1}. \quad \text{(A1)}$$

It is straightforward to show that for $\rho<1$, $g_{NN}^* = ((1-\rho)/N)[(N/g)^\rho - 1] < 0$, and $g_{NI}^* = \gamma((1-\rho)/I)(I/g)^\rho > 0$, while with $\rho=1$ (perfect substitutes), $g_{NN}^*=g_{NI}^*=0$. Note that Ng_{NN}^* and Ig_{NI}^* are elasticities of complementarity, as in Lalonde and Topel (1991). These terms sum to zero with constant returns and capture the complementarity that makes immigration effects ambiguous.

To produce the expression in the text, note that the native labor demand curve (ignoring firing costs) is defined by

$$w_N = f'\theta g_N.$$

Differentiate with respect to N and re-arrange to show that along the labor demand curve

$$(\partial w_N/\partial N)(N/w_N) \equiv \xi_{NN} = N(f''/f')\theta g_N + Ng_{NN}^*.$$

Similarly,

$$(\partial w_N/\partial I)(I/w_N) \equiv \xi_{NI} = I(f''/f')\theta g_I + Ig_{NI}^*.$$

Finally, substitute into (A1) to get

$$\partial N/\partial I = [\xi_{NI}/I][(1/\varepsilon)N^{-1} - \xi_{NN}N^{-1}]^{-1}.$$

Derivation of long-run employment change, (6)

Note that

$$\partial N^*/\partial M = m(\partial N/\partial M) + N(\partial m/\partial M).$$

Write the first order condition, (2a), as

$$\mathit{ln}f'(\theta g[N,Mm^{-1}]) + \mathit{ln}\theta + \mathit{ln}g_N[N,Mm^{-1}] = (1/\varepsilon)\mathit{ln}N + (1/\varepsilon)\mathit{ln}m + \mathit{ln}(1 + \varphi\lambda c_N) - \mathit{ln}P - \mathit{ln}(1-r).$$

Differentiate and re-arrange to get

$$(\partial N/\partial M)\{(1/\varepsilon)N^{-1} - (f''/f')\theta g_N - g_{NN}^*\} = m^{-1}\{(f''/f')\theta g_I + g_{NI}^*\}\{1-(\partial \mathit{ln}m/\partial \mathit{ln}M)\} - (1/\varepsilon)m^{-1}(\partial m/\partial M),$$

so

$$\begin{aligned} m(\partial N/\partial M) &= \{(f''/f')\theta g_I + g_{NI}^*\}\{(1/\varepsilon)N^{-1} - (f''/f')\theta g_N - g_{NN}^*\}^{-1}\{1-(d\mathit{ln}m/d\mathit{ln}M)\} - (\varepsilon\Delta)^{-1}(\partial m/\partial M) \\ &= (N/I)e(N, \varepsilon)\{1-(\partial \mathit{ln}m/\partial \mathit{ln}M)\} - (\varepsilon\Delta)^{-1}(\partial m/\partial M). \end{aligned}$$

Now, add $N(\partial m/\partial M)$ to get

$$\partial N^*/\partial M = (N/I)e(N, \varepsilon)\{1-(\partial \mathit{ln}m/\partial \mathit{ln}M)\} + (\partial m/\partial M)[N - (\varepsilon\Delta)^{-1}].$$

Note that $N - (\varepsilon\Delta)^{-1} = [N\Delta - (1/\varepsilon)]/\Delta$ and $N\Delta = (1/\varepsilon) - \xi_{NN}$, so

$$\partial N^*/\partial M = (N/I)e(N, \varepsilon)\{1-(\partial lnm/\partial lnM)\} - (\partial m/\partial M)[\xi_{NN}/\Delta].$$

Multiply both sides by $I/N = M/mN$ to get

$$(\partial N^*/\partial M)(M/N^*) = e(N, \varepsilon)\{1-(\partial lnm/\partial lnM)\} - \xi_{NN}[\varepsilon^{-1} - \xi_{NN}]^{-1} (\partial lnm/\partial lnM).$$

Effect of immigration on profits in the short run

As noted in the text, profits clearly increase if the wages of natives fall after immigration. But suppose immigrant-native complementarity is strong enough that wages of natives increase. With m taken to be fixed, solve for $\partial w_N/\partial M$ and $\partial w_I/\partial M$ by differentiating the first order conditions, (2a) and (2b), then substitute into $\partial \Pi/\partial M$ to get:

$$-(N(\partial w_N/\partial M) + I(\partial w_I/\partial M)) = -[g_N (\partial N/\partial M) + (g_I/m)]f''\theta^2 g.$$

If native wages go up, then native employment must go up as well, so $\partial N/\partial M$ is positive. The expression in brackets is therefore also positive and profits increase. If the production function exhibits constant returns, however, $f''(\cdot)=0$, and profits are unchanged.

Proof that the long-run impact is zero with ρ=1

Again, solve for $\partial w_N/\partial M$ and $\partial w_I/\partial M$ by differentiating the first order conditions, (2a) and (2b), but retain terms involving $\partial lnm/\partial lnM$. Then substitute the resulting expressions into $\partial \Pi/\partial M = 0$, or, equivalently, $N(\partial w_N/\partial M) + I(\partial w_I/\partial M)=0$, to obtain

$$\partial lnm/\partial lnM = 1 + m(\partial N/\partial M)(g_N/g_I).$$

Next, use this to substitute for $\partial lnm/\partial lnM$ in the expression for $(\partial N^*/\partial M)(M/N^*)$. Finally, re-arrange, and use the fact that if $\rho=1$, then $\xi_{NI} N g_N = \xi_{NN} I g_I$ to show that $(\partial N^*/\partial M)(M/N^*)=0$.

REFERENCES

Acemoglu, Daron, and Joshua Angrist (2001), "Consequences of Employment Protection? The Case of the Americans with Disabilities Act," *Journal of Political Economy* 109, 915-957.

Altonji, Joseph G. and David Card (1991), "The Effects of Immigration on the Labor Market Outcomes of Less-skilled Natives." In John M. Abowd and Richard B. Freeman, editors, *Immigration, Trade and Labor*. Chicago: University of Chicago Press.

Bauer, Thomas, and Klaus F. Zimmermann (1999), "Assessment of Possible Migration Pressure and its Labour Market Impact Following EU Enlargement to Central and Eastern Europe," IZA Research Report 3, Bonn: The IZA, July.

__________, Magnus Lofstrom, and Klauss F. Zimmermann (2000), "Immigration Policy Assimilation of Immigrants and Native's Sentiment Towards Immigrants: Evidence from 12 OECD Countries," IZA Discussion Paper No. 187, Bonn: The IZA, August.

Bentolila, Samuel, and Giuseppe Bertola (1990), 'Firing Costs and Labour Demand: How Bad is Eurosclerosis?," *Review of Economic Studies* 57, 381-402.

Bertrand, Marianne, and Francis Kramarz (2001), "Does Entry Regulation Hinder Job Creation? Evidence from the French Retail Industry," NBER Working Paper No. 8211, April.

Blanchard, Olivier, and Justin Wolfers (2000), "The Role of Shocks and Institutions in the Rise of European Unemployment: The Aggregate Evidence,"*Economic Journal* 110, C1-33.

Borjas, George J. (1994), "The Economics of Immigration," *Journal of Economic Literature* 32, 1667-717.

__________ (1995), "The Economic Benefits from Immigration," *Journal of Economic Perspectives* 9, 3-22.

Card, David E. (1990), "The Impact of the Mariel Boatlift on the Miami Labor Market." *Industrial and Labor Relations Review* 43, 245-257.

__________ (2001), "Immigrant Inflows, Native Outflows, and the Local Labor Market Impacts of Higher Immigration," *Journal of Labor Economics* 19, 22-64.

Carrington, William, and Pedro deLima (1996), "The Impact of 1970s Repatriates from Africa on the Portuguese Labor Market," *Industrial and Labor Relations Review* 49, 330-47.

Dolado, Juan J., R. Duce, and J.F. Jimeno (1996), "The Effects of Migration on the Relative Demand of Skilled Versus Unskilled Labour: Evidence from Spain," CEPR Discussion Paper No. 1476, September.

Eurostat (1998), *Labor Force Survey: Methods and Definitions,* Brussels: The European Commission.

Friedberg, Rachel M. and Jennifer Hunt (1995). "The Impact of Immigrants on Host Country Wages, Employment, and Growth." *Journal of Economic Perspectives* 9, 23-44

Hopenhayn, Hugo, and Richard Rogerson (1993), “Job Turnover and Policy Evaluation: A General Equilibrium Analysis, *Journal of Political Economy* 101, 915-938.

Hunt, Jennifer (1992). “The Impact of the 1962 Repatriates from Algeria on the French Labor Market,” *Industrial and Labor Relations Review* 45, 556-72.

Ichino, Andrea (1993), “The Economic Impact of Immigration on the Host Country,” in G. Luciani, ed., *Migration Policies in Europe and the United States*, Kluwer: Dordrecht, 145-162.

Ichino, Andrea, and Regina Riphahn (2000), “Measuring the Effect of Employment Protection on Shirking,” Florence: European University Institute, mimeo, November.

Lalonde, Robert J. and Robert H. Topel (1991). “Labor Market Adjustments to Increased Immigration.” In John M. Abowd and Richard B. Freeman, editors, *Immigration, Trade and Labor*. Chicago: University of Chicago Press.

Johnson, G. (1980). “The Labor Market Effects of Immigration,” *Industrial and Labor Relations Review* 33, 331-341.

Nicoletti, Giuseppe, Stefano Scarpetta, and Olivier Boylaud (2000), “Summary Indicators of Product Market Regulation with an Extension to Employment Protection Legislation,” OECD Economics Department Working Paper No. 226, Paris: The OECD.

Nickell, S. (1997), “Unemployment and Labor Market Rigidities: Europe versus North America,” *Journal of Economic Perspectives* 11, 55-74.

OECD (1994), “*The OECD Jobs Study: Facts, Analysis, Strategies*, Paris: The Organization for Economic Cooperation and Development.

OECD (1995), *Trends in International Migration, Annual Report 1994, 1995 Edition*, Paris: The Organization for Economic Cooperation and Development.

OECD (1999a), *Trends in International Migration, Annual Report, 1999 Edition*, Paris: The Organization for Economic Cooperation and Development.

OECD (1999b), *OECD Employment Outlook*, 1999, Paris: The Organization for Economic Cooperation and Development, June.

OECD (2000), *OECD Employment Outlook*, 2000, Paris: The Organization for Economic Cooperation and Development, June.

Pischke, J.S., and J. Velling (1997), “Employment Effects of Immigration to Germany: An Analysis Based on Local Labor Markets,” *Review of Economics and Statistics* 79, 594-604.

Razin, Assaf, and Efraim Sadka (1996), “Suppressing Resistance to Low-Skill Migration,” *International Tax and Public Finance* 3, 413-424.

________ and _________ (1995), "Resisting Migration: Wage Rigidity and Income Distribution," *American Economic Review* 85(2), 312-316.

Rodrik, Dani (1997), *Has Globalization Gone Too Far?*, Washington, DC: The Institute for International Economics.

Saint-Paul, Gilles (1996), *Dual Labor Markets: A Macroeconomic Perspective*, Cambridge: The MIT Press.

Schmidt, Christoph M., Anette Stilz, and Klauss F. Zimmermann (1994), "Mass Migration, Unions, and Government Intervention," *Journal of Public Economics* 55, 185-201.

Schwartz, Aba (1973), "Interpreting the Effect of Distance on Migration," *The Journal of Political Economy* 81, 1153-1169.

Winter-Ebmer, Rudolf, and J. Zweimuller (1997), "Immigration, Trade, and Austrian Unemployment," in M. Landesmann and E. Streissler, eds., *Unemployment in Europe*, London: MacMillan.

Table 1
Descriptive Statistics

Country	LFS Coverage	Labor Force in 1999 (1000s)	Proportion Non-EU Non-Nationals			Proportion EU Non-Nationals			E/P (Nationals)		
			1989	1996	1999	1989	1996	1999	1989	1996	1999
A. EU-12											
FRG	1983-90	–	5.01	–	–	3.22	–	–	71.6	–	–
Germany	1991-99	36,622	–	6.08	6.13	–	2.79	2.73	–	73.6	75.7
France	1983-99	24,992	3.72	3.88	3.85	2.94	2.40	2.30	72.2	71.9	72.4
Italy	1992-99	21,892	–	0.38	0.86	–	0.06	0.13	–	58.6	60.3
Netherlands	83,85,87-99	7,183	2.19	2.09	1.74	1.67	1.74	1.74	66.3	72.8	78.1
Belgium	1983-99	4,206	2.28	2.27	2.66	4.84	5.96	6.12	64.9	74.3	71.0
Luxembourg	1983-99	174	2.26	3.98	4.68	30.0	35.8	37.2	64.0	65.3	68.6
U.K.	1983-99	25,829	2.94	2.18	2.39	1.77	1.55	1.68	76.5	75.3	77.5
Ireland	1983-99	1,502	0.47	0.80	1.00	2.16	2.98	2.66	56.6	63.6	70.2
Denmark	1983-99	2,564	1.16	1.40	1.59	0.79	0.79	1.00	81.3	79.8	82.5
Greece	1983-99	4,034	0.51	1.67	3.78	0.11	0.19	0.21	64.2	64.6	64.7
Spain	1986-99	15,118	0.15	0.48	0.89	0.14	0.33	0.37	55.6	54.9	60.4
Portugal	1986-99	4,375	0.59	1.03	1.06	–	0.22	0.34	72.9	72.0	75.8
B. EU – Other											
Austria	1995-99	3,590	–	8.47	8.19	–	1.22	1.44	–	74.3	75.6
Finland	1995-99	2,406	–	0.73	1.09	–	–	0.14	–	69.0	76.1
Sweden	1995-99	4,007	–	3.89	3.01	–	0.71	1.28	–	78.8	79.5
C. Other EEA											
Iceland	1995-99	128	–	–	1.27	–	–	–	–	87.1	89.0
Norway	1995-99	2,072	–	1.82	1.97	–	1.00	1.18	–	81.0	83.5
Switzerland	1996-99	3,479	–	5.56	5.83	–	17.5	16.0	–	82.7	85.1

Notes: The table reports weighted counts in thousands and proportions in the indicated period. All statistics are for men and women aged 20-59 in the Eurostat Labor Force Survey.

Table 2
Nationality and Nativity in 1996

Country	Proportion of Non-nationals who are Foreign Born		Naturalization Rate			
			All Foreign Born		Recent Arrivals	
	EU Nationals (1)	Non-EU Nationals (2)	EU Immigrants (3)	Non-EU Immigrants (4)	EU Immigrants (5)	Non-EU Immigrants (6)
A. EU-12						
Germany	72.3	95.1	12.9	66.0	10.0	48.1
France	93.8	96.0	37.4	50.6	0.0	0.3
Netherlands	81.2	97.2	36.7	58.0	4.2	14.5
Belgium	68.4	93.3	30.3	35.6	2.7	8.9
Luxembourg	87.0	96.8	7.8	13.9	1.1	0.5
U.K.	94.2	98.1	28.0	57.4	1.1	7.5
Ireland	94.0	94.1	46.3	29.1	12.0	10.5
Denmark	89.6	96.4	43.1	46.1	6.2	7.7
Greece	82.0	95.4	81.3	60.2	57.5	43.5
Spain	83.2	97.7	70.1	61.8	0.0	0.0
Portugal	74.2	98.1	77.7	94.1	56.4	31.3
B. EU – Other						
Austria	69.0	93.9	44.3	27.2	6.7	5.5
Finland	59.2	98.6	83.0	25.7	0.0	0.0
Sweden	69.4	92.1	42.5	56.5	22.4	10.2
C. Other EEA						
Iceland	82.6	84.1	58.6	44.0	0.0	0.0
Norway	827	94.3	41.5	45.5	1.9	2.8

Note: Statistics are for men and women aged 20-59. Data for Germany are for 1992. The LFS does not distinguish nationality and nativity for Italy. Entries of zero denote small counts in cells or cells below the LFS disclosure threshold. Nativity variables are missing for Switzerland.

Table 3
Characteristics of Immigrants and Natives

	Men					Women				
		Non-EU		Yugoslavs			Non-EU		Yugoslavs	
	Natives	Recent	Veteran	Recent	Veteran	Natives	Recent	Veteran	Recent	Veteran
Country	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
A. Labor Force Participation Rate										
Germany	90.3	77.5	88.8	–	–	75.4	50.4	68.7	–	–
France	88.2	67.7	87.5	72.2	86.1	73.9	33.4	57.8	69.6	72.1
Italy	82.8	93.5	90.4	75.2	69.0	51.0	51.9	57.3	39.4	–
Netherlands	90.2	55.8	78.3	50.4	75.5	68.1	37.7	53.9	32.3	60.0
Belgium	85.3	72.2	76.1	78.1	79.1	65.3	33.1	43.7	21.7	52.2
Luxembourg	85.1	86.7	84.8	88.3	84.7	51.4	42.0	53.9	45.1	44.4
U.K.	90.0	71.1	86.7	59.6	80.7	73.6	45.1	61.7	52.6	58.5
Ireland	87.7	57.4	84.7	–	–	57.6	34.8	61.0	–	–
Denmark	91.0	68.0	76.1	64.7	90.0	80.6	34.1	65.0	21.0	71.9
Greece	89.0	93.3	92.4	93.7	97.5	53.8	67.7	60.3	71.1	54.8
Spain	86.3	94.1	90.1	–	81.0	54.2	62.7	61.5	–	56.6
Portugal	89.2	89.6	74.5	–	–	71.4	68.8	59.6	–	–
Austria	88.2	86.8	90.9	90.9	92.2	69.7	52.8	69.6	56.5	78.7
Finland	84.8	76.6	80.5	–	61.3	79.5	53.4	69.1	–	–
Sweden	88.7	61.9	81.1	65.2	79.2	84.4	41.7	68.1	49.6	63.9
Norway	89.8	75.5	76.3	46.7	80.4	80.8	46.9	64.4	34.0	69.1
B. Employment/Population										
Germany	83.3	62.2	80.4	–	–	68.7	39.5	61.5	–	–
France	80.0	42.1	69.8	45.4	67.5	64.0	20.8	43.1	66.1	55.6
Italy	75.3	87.0	85.4	58.7	69.0	42.9	39.6	50.0	20.8	–
Netherlands	87.4	37.3	68.1	32.7	65.9	64.2	26.9	48.1	19.1	52.8
Belgium	80.0	51.6	58.7	27.4	64.1	58.3	21.0	32.2	21.7	35.3
Luxembourg	83.8	80.5	80.0	79.3	79.8	50.0	37.9	49.4	39.3	37.5
U.K.	83.1	60.2	76.8	28.3	56.1	69.7	37.7	56.2	41.1	55.5
Ireland	78.9	48.9	77.7	–	–	52.2	29.1	55.8	–	–
Denmark	87.2	52.8	64.4	45.7	74.9	75.3	27.5	54.0	21.0	61.8
Greece	83.5	82.3	83.4	83.6	86.6	45.7	51.6	49.0	48.8	49.6
Spain	73.6	73.9	75.7	84.3	81.0	39.7	51.1	44.7	–	56.6
Portugal	86.2	76.7	60.4	–	–	68.1	53.2	47.9	–	–
Austria	84.4	76.9	82.0	78.3	83.0	66.7	44.9	64.3	48.9	74.2
Finland	74.0	48.6	61.6	–	46.3	69.4	38.0	55.2	–	–
Sweden	81.4	38.9	60.9	37.9	62.1	78.8	23.8	55.3	25.8	51.2
Norway	86.9	64.2	69.0	31.9	80.4	77.8	38.0	59.3	21.4	52.3

Notes: The table reports statistics for male and female immigrants and natives aged 20-59 in 1995-9. Data are unavailable for Switzerland and unreliable for Iceland. A dash denotes entries below the LFS confidentiality threshold.

Table 4
OLS Estimates of Effects of Non-EU Share 92-99

Period	Pooled (1)	By Age Group: Under 40 (2)	By Age Group: Over 40 (3)	By Schooling Group: Low (4)	By Schooling Group: Medium (5)	By Schooling Group: High (6)
A. Men						
Non-Nationals	-0.011	-0.015	-0.008	-0.002	0.001	-0.033
	(0.021)	(0.029)	(0.022)	(0.024)	(0.031)	(0.027)
N	642	321	321	214	214	214
Foreign Born	0.012	0.003	0.021	0.022	0.002	0.011
	(0.019)	(0.030)	(0.016)	(0.017)	(0.036)	(0.011)
N	592	296	296	196	198	198
Recent Arrivals	-0.011	-0.010	-0.013	-0.030	0.006	-0.013
	(0.011)	(0.014)	(0.009)	(0.014)	(0.018)	(0.011)
N	586	293	293	194	196	196
B. Women						
Non-Nationals	-0.029	-0.010	-0.047	-0.010	0.006	-0.082
	(0.040)	(0.039)	(0.057)	(0.055)	(0.056)	(0.044)
N	642	321	321	214	214	214
Foreign Born	0.022	0.026	0.019	0.041	0.018	0.004
	(0.028)	(0.034)	(0.031)	(0.042)	(0.050)	(0.018)
N	592	296	296	196	198	198
Recent Arrivals	-0.030	-0.025	-0.035	-0.056	-0.010	-0.031
	(0.020)	(0.019)	(0.026)	(0.032)	(0.029)	(0.017)
N	586	293	293	194	196	196

Notes: The table reports estimates of the coefficient on log share of non-nationals or non-natives from non-EU countries in equation (7) in the text. All models include country and year effects. Pooled models include main effects for demographic groups. Robust standard errors are reported in parenthesis. The estimation sample for this table excludes Italy.

Table 5
OLS Estimates by Sex, Age, and Schooling Group

Period and controls	Non-National share	Pooled (1)	By Age Group: Under 40 (2)	By Age Group: Over 40 (3)	By Schooling Group: Low (4)	By Schooling Group: Medium (5)	By Schooling Group: High (6)
A. Men							
1992-99	Non-EU	-0.013	-0.015	-0.011	-0.008	0.003	-0.035
		(0.021)	(0.028)	(0.022)	(0.022)	(0.029)	(0.028)
	EU	0.034	0.045	0.023	0.056	0.015	0.030
		(0.023)	(0.032)	(0.018)	(0.018)	(0.040)	(0.024)
N		684	342	342	228	228	228
1983-99	Non-EU	-0.021	-0.037	-0.004			
		(0.008)	(0.008)	(0.005)			
	EU	0.036	0.053	0.018			
		(0.016)	(0.014)	(0.010)			
N		402	201	201			
1983-99	Non-EU	-0.011	-0.012	-0.010			
Trends		(0.015)	(0.028)	(0.007)			
	EU	0.022	0.028	0.016			
		(0.019)	(0.093)	(0.006)			
N		402	201	201			
B. Women							
1992-99	Non-EU	-0.033	-0.016	-0.051	-0.027	0.004	-0.077
		(0.038)	(0.038)	(0.056)	(0.051)	(0.053)	(0.043)
	EU	0.047	0.053	0.040	0.083	0.023	0.034
		(0.041)	(0.038)	(0.060)	(0.051)	(0.052)	(0.034)
N		684	342	342	228	228	228
1983-99	Non-EU	-0.026	-0.026	-0.026			
		(0.026)	(0.016)	(0.026)			
	EU	0.086	0.092	0.081			
		(0.041)	(0.026)	(0.031)			
N		402	201	201			
1983-99	Non-EU	-0.012	-0.028	-0.002			
Trends		(0.048)	(0.012)	(0.015)			
	EU	0.008	0.018	-0.002			
		(0.049)	(0.011)	(0.013)			
N		402	201	201			

Notes: The table reports estimates of the coefficient on log share of non-nationals from non-EU and EU countries in equation (7) in the text. All models include country and year effects. Pooled models include main effects for demographic groups. Robust standard errors are reported in parenthesis.

Table 6
First Stage Estimates

Instruments	Distance to Nearest Big City				Distance to Capitals			
	No Trends		With Trends		No Trends		With Trends	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Non-Nationals from Non-EU Countries								
Bosnia War (1991-95) × Distance from Sarajevo	-0.830 (0.182)	-0.763 (0.224)	-0.624 (0.234)	-0.601 (0.228)	-0.754 (0.256)	-0.515 (0.336)	-0.543 (0.273)	-0.509 (0.261)
No War (1996-97) × Distance from Sarajevo	-0.712 (0.222)	-0.647 (0.254)	-0.743 (0.296)	-0.686 (0.300)	-0.556 (0.293)	-0.320 (0.359)	-0.654 (0.349)	-0.577 (0.348)
Kosovo War (1998-99) × Distance from Pristina	-0.924 (0.266)	-0.865 (0.300)	-1.082 (0.296)	-1.010 (0.306)	-0.820 (0.369)	-0.598 (0.436)	-1.089 (0.352)	-0.990 (0.355)
Prewar (1988-90) × Distance from Sarajevo		-0.090 (0.186)		0.096 (0.159)		-0.316 (0.209)		0.121 (0.158)
F-Statistic for excluded instruments	6.96	3.97	6.67	5.39	3.18	1.25	5.98	4.79
B. Non-Nationals from EU Countries								
Bosnia War × Distance from Sarajevo	-0.122 (0.103)	-0.181 (0.136)	-0.170 (0.181)	-0.168 (0.182)	-0.099 (0.137)	-0.126 (0.180)	-0.093 (0.041)	-0.009 (0.193)
No War × Distance from Sarajevo	0.158 (0.149)	0.100 (0.170)	-0.098 (0.267)	-0.094 (0.284)	0.232 (0.186)	0.205 (0.218)	0.081 (0.323)	0.108 (0.336)
Kosovo War × Distance from Pristina	0.104 (0.120)	0.051 (0.143)	-0.219 (0.290)	-0.214 (0.322)	0.189 (0.150)	0.164 (0.184)	-0.037 (0.316)	-0.002 (0.340)
Prewar (1988-90) × Distance from Sarajevo		0.081 (0.115)		0.007 (0.165)		0.035 (0.118)		0.042 (0.160)
F-Statistic for excluded instruments	1.27	1.34	0.43	0.43	1.34	1.28	0.14	0.14

Note: The table reports the effects of dummies for Balkan war years interacted with the distance between Sarajevo (or Pristina) and the nearest big cities (or entry ports) on the log share of non-nationals from non-EU and EU countries. This is equation (8) in the text. Coefficients are scaled to show the impact of 1,000 miles distance. The Bosnia war years are 1991-95 and the Kosovo war years are 1998-99. The years 1996-97 are the inter-war period. Robust standard errors are reported in parenthesis. All models include country and year effects. Sample sizes are 201 observations for non-nationals from non-EU countries and 202 observations for non-nationals from EU countries.

Table 7: IV Estimates

Non-national share	Sex	No Trends			With Trends		
			By Age Group			By Age Group	
		Pooled (1)	Under 40 (2)	Over 40 (3)	Pooled (4)	Under 40 (5)	Over 40 (6)
A. Without share of Non-Nationals from EU Countries							
Non-EU	Men	-0.050	-0.082	-0.018	-0.034	-0.028	-0.040
		(0.023)	(0.030)	(0.016)	(0.029)	(0.027)	(0.018)
Non-EU	Women	-0.245	-0.189	-0.301	-0.058	-0.034	-0.082
		(0.093)	(0.070)	(0.102)	(0.112)	(0.030)	(0.046)
N		420	210	210	420	210	210
B. With share of Non-Nationals from EU Countries							
Non-EU	Men	-0.050	-0.089	-0.011	-0.043	-0.042	-0.045
		(0.016)	(0.020)	(0.011)	(0.034)	(0.031)	(0.020)
EU		0.059	0.094	0.024	0.032	0.037	0.027
		(0.016)	(0.020)	(0.012)	(0.017)	(0.016)	(0.010)
Non-EU	Women	-0.210	-0.166	-0.253	-0.067	-0.046	-0.088
		(0.064)	(0.043)	(0.062)	(0.132)	(0.033)	(0.053)
EU		0.232	0.203	0.260	0.026	0.025	0.026
		(0.066)	(0.044)	(0.064)	(0.067)	(0.017)	(0.027)
N		402	201	201	402	201	201

Notes: The table reports IV estimates of equation (7) in the text. The non-EU immigration-share variable is treated as endogenous, while the EU share variable is treated as exogenous.

Table 8
Interactions with Institutions: Estimates for Men

		OLS			2SLS		
			By Age Group			By Age Group	
		Pooled	Under 40	Over 40	Pooled	Under 40	Over 40
Interaction	Regressor	(1)	(2)	(3)	(4)	(5)	(6)
		A. Institutions one at a time					
Labor Standards	Main Effect	-0.0270	-0.0445	-0.0096	-0.0314	-0.0709	0.0081
		(0.0102)	(0.0120)	(0.0075)	(0.0252)	(0.0343)	(0.0208)
	Interaction	-0.0077	-0.0099	-0.0054	-0.0361	-0.0465	-0.0256
		(0.0034)	(0.0040)	(0.0025)	(0.0127)	(0.0174)	(0.0105)
N		334	167	167	334	167	167
Replacement Rate	Main Effect	-0.0274	-0.0447	-0.0102	0.0511	0.1083	-0.0061
		(0.0102)	(0.0121)	(0.0074)	(0.0557)	(0.0892)	(0.0404)
	Interaction	-0.0009	-0.0011	-0.0008	0.0004	0.0000	0.0008
		(0.0004)	(0.0005)	(0.0003)	(0.0010)	(0.0015)	(0.0007)
N		334	167	167	334	167	167
Barriers to Entrepreneurship	Main Effect	-0.0270	-0.0439	-0.0101	-0.0609	-0.1165	-0.0052
		(0.0079)	(0.0094)	(0.0058)	(0.0199)	(0.0278)	(0.0167)
	Interaction	-0.0308	-0.0381	-0.0235	-0.0647	-0.0335	-0.0959
		(0.0148)	(0.0175)	(0.0109)	(0.0370)	(0.0517)	(0.0311)
N		368	184	184	368	184	184
		B. Institutions Together					
Standards and Replacement Rate	Main Effect	-0.0284	-0.0460	-0.0108	-0.0322	-0.0799	0.0155
		(0.0102)	(0.0119)	(0.0074)	(0.0285)	(0.0422)	(0.0208)
	Labor Standards	-0.0061	-0.0082	-0.0041	-0.0284	-0.0465	-0.0104
		(0.0035)	(0.0041)	(0.0025)	(0.0103)	(0.0152)	(0.0075)
	Replacement Rate	-0.0007	-0.0008	-0.0007	-0.0009	-0.0017	-0.0001
		(0.0004)	(0.0005)	(0.0003)	(0.0010)	(0.0014)	(0.0007)
N		334	167	167	334	167	167
Standards, Rep. Rate, and Barriers	Main Effect	-0.0347	-0.0526	-0.0168	-0.0552	-0.1024	-0.0080
		(0.0117)	(0.0137)	(0.0084)	(0.0331)	(0.0440)	(0.0256)
	Labor Standards	-0.0010	-0.0028	0.0007	0.0124	0.0146	0.0103
		(0.0058)	(0.0068)	(0.0042)	(0.0101)	(0.0134)	(0.0078)
	Replacement Rate	-0.0009	-0.0010	-0.0008	0.0003	0.0003	0.0004
		(0.0005)	(0.0005)	(0.0003)	(0.0008)	(0.0010)	(0.0006)
	Barriers	-0.0280	-0.0295	-0.0265	-0.1453	-0.2044	-0.0862
		(0.0255)	(0.0300)	(0.0185)	(0.0537)	(0.0713)	(0.0416)
N		334	167	167	334	167	167

Note: The table reports estimates of main effects and interaction terms in equation (9) in the text. All models include the EU foreign share as an exogenous regressor. Instruments for the foreign share are as in Table 7, plus interactions with institutional measures. Main effects are evaluated at the median institution (5 for labor standards, 63 for replacement rates, and 1.72 for barriers to entrepreneurship).

Table 9
Interactions with Institutions: Estimates for Men with Country Trends

		OLS			2SLS		
			By Age Group			By Age Group	
		Pooled	Under 40	Over 40	Pooled	Under 40	Over 40
Interaction	Regressor	(1)	(2)	(3)	(4)	(5)	(6)
		A. Institutions one at a time					
Labor Standards	Main Effect	-0.0144	-0.0189	-0.0100	-0.0496	-0.0674	-0.0319
		(0.0157)	(0.0144)	(0.0089)	(0.0643)	(0.0642)	(0.0375)
	Interaction	-0.0016	-0.0032	-0.0001	-0.0178	-0.0259	-0.0093
		(0.0056)	(0.0052)	(0.0032)	(0.0189)	(0.0189)	(0.0110)
N		334	167	167	334	167	167
Replacement Rate	Main Effect	-0.0224	-0.0272	-0.0176	-0.0448	-0.0033	-0.0862
		(0.0159)	(0.0143)	(0.0086)	(0.0656)	(0.0618)	(0.0439)
	Interaction	-0.0015	-0.0016	-0.0013	-0.0009	0.0005	-0.0022
		(0.0007)	(0.0006)	(0.0004)	(0.0016)	(0.0015)	(0.0011)
N		334	167	167	334	167	167
Barriers to Entrepreneurship	Main Effect	-0.0134	-0.0150	-0.0118	-0.0491	-0.0481	-0.0500
		(0.0126)	(0.0117)	(0.0073)	(0.0367)	(0.0398)	(0.0259)
	Interaction	-0.0039	-0.0045	-0.0033	-0.1265	-0.1549	-0.0980
		(0.0195)	(0.0181)	(0.0113)	(0.0540)	(0.0586)	(0.0382)
N		368	184	184	368	184	184
		B. Institutions Together					
Standards and Replacement Rate	Main Effect	-0.0241	-0.0298	-0.0184	-0.0700	-0.0493	-0.0906
		(0.0161)	(0.0145)	(0.0087)	(0.0674)	(0.0618)	(0.0475)
	Labor Standards	-0.0034	-0.0051	-0.0016	-0.0181	-0.0193	-0.0169
		(0.0056)	(0.0051)	(0.0031)	(0.0214)	(0.0196)	(0.0151)
	Replacement Rate	-0.0015	-0.0017	-0.0013	-0.0023	-0.0013	-0.0032
		(0.0007)	(0.0006)	(0.0004)	(0.0020)	(0.0018)	(0.0014)
N		334	167	167	334	167	167
Standards, Rep. Rate, and Barriers	Main Effect	-0.0238	-0.0299	-0.0177	-0.0460	-0.0186	-0.0734
		(0.0162)	(0.0146)	(0.0088)	(0.0526)	(0.0505)	(0.0368)
	Barriers	-0.0111	0.0026	-0.0247	-0.1532	-0.1462	-0.1601
		(0.0384)	(0.0346)	(0.0207)	(0.0976)	(0.0937)	(0.0682)
	Labor Standards	-0.0009	-0.0057	0.0039	0.0287	0.0259	0.0316
		(0.0102)	(0.0092)	(0.0055)	(0.0320)	(0.0307)	(0.0224)
	Replacement Rate	-0.0016	-0.0017	-0.0014	-0.0020	-0.0012	-0.0027
		(0.0007)	(0.0006)	(0.0004)	(0.0016)	(0.0015)	(0.0011)
N		334	167	167	334	167	167

Note: The table reports estimates of main effects and interaction terms in equation (9) in the text. All models include the EU foreign share as an exogenous regressor. Instruments for the foreign share are as in Table 7, plus interactions with institutional measures. Main effects are evaluated at the median institution (5 for labor standards, 63 for replacement rates, and 1.72 for barriers to entrepreneurship).

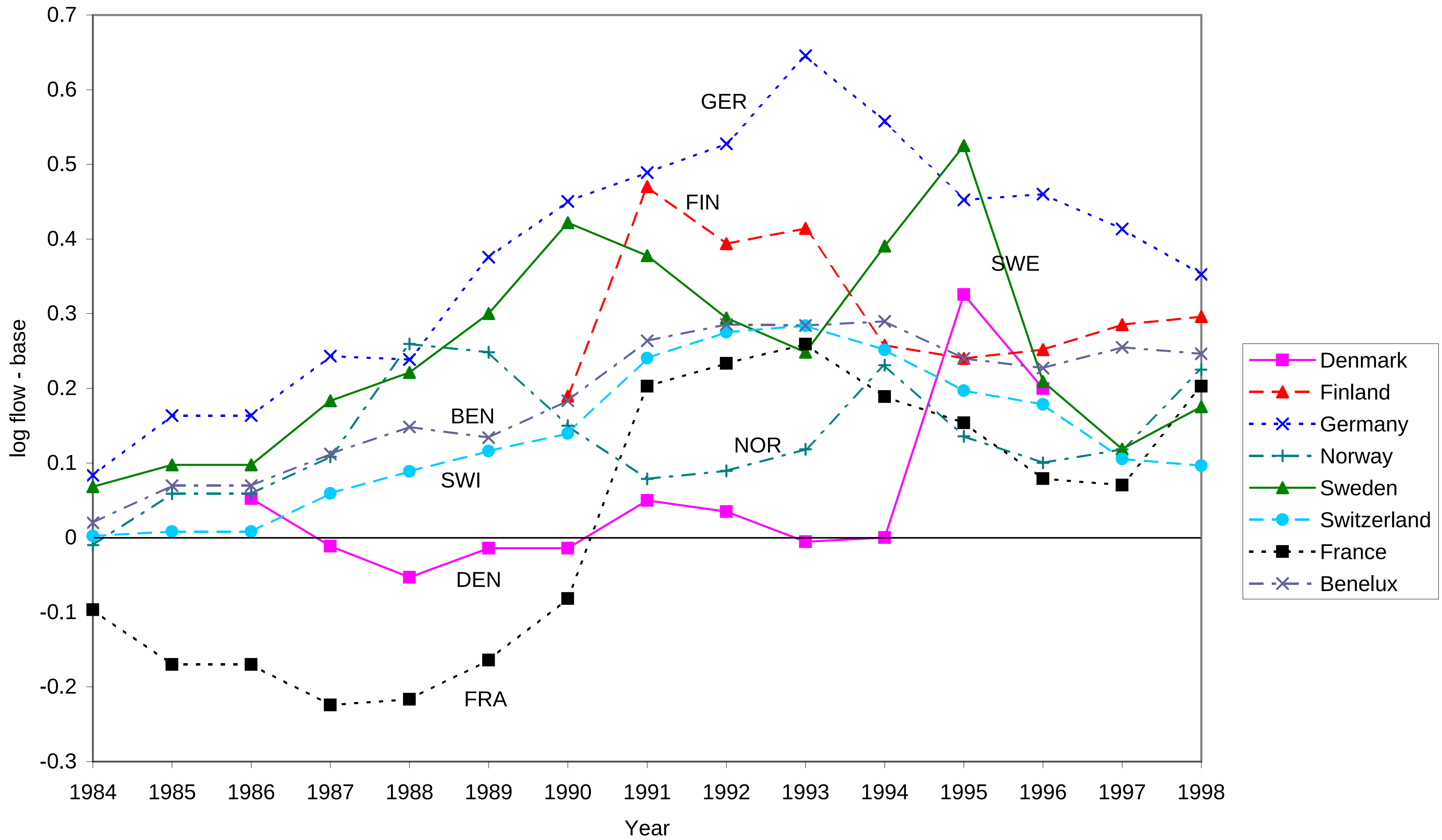

Figure 1. Foreign inflows. Log count minus 1983 base. Data from population registers except for France.
Source: OECD Trends in International Migration

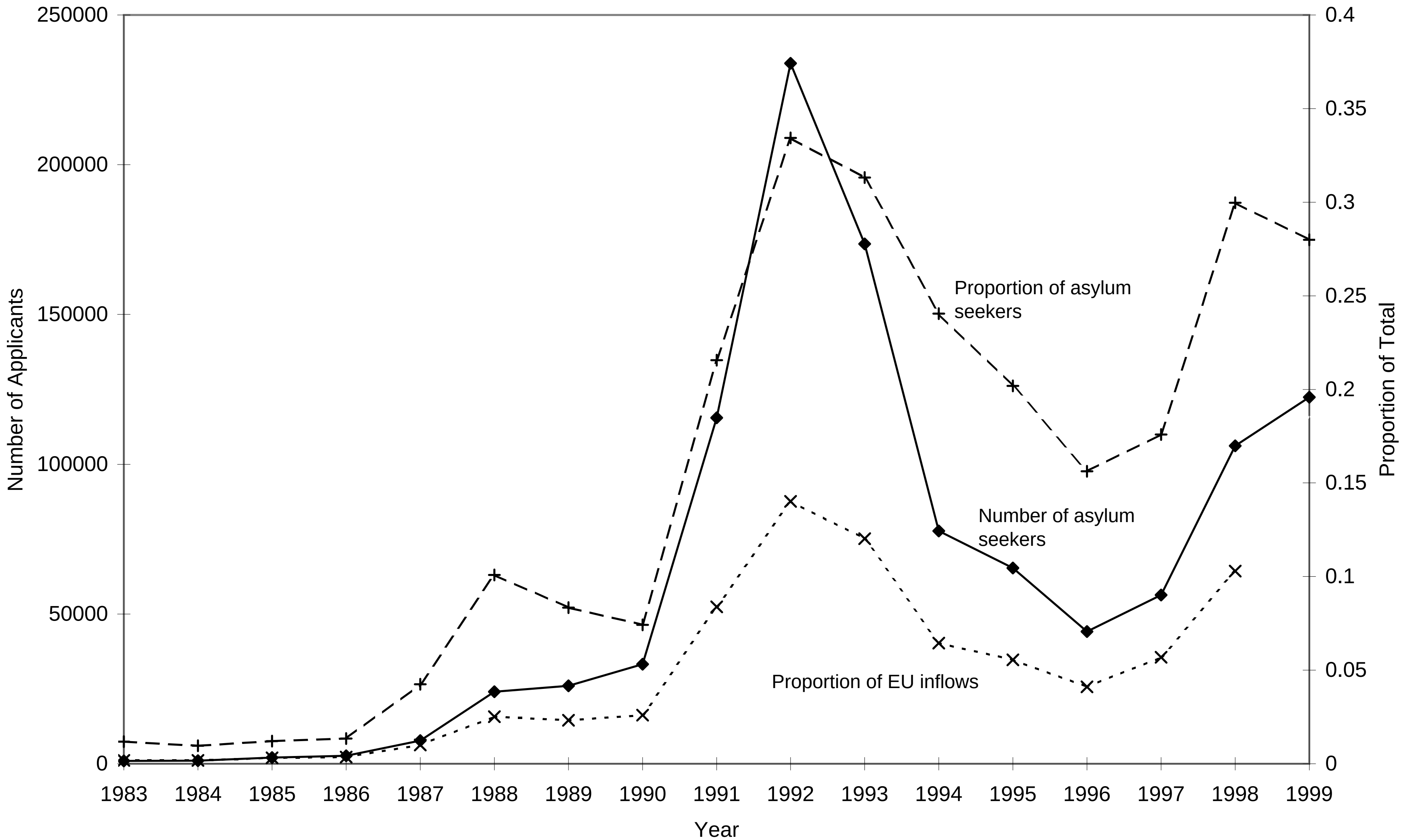

Figure 2. Number of Yugoslavian asylum applicants to Europe, Yugoslav proportion of total, and Yugoslavians as a proportion of total foreign inflows to selected EU countries. Data for 1992-99 include asylum applicants from both Yugoslavia and Bosnia. Sources: Refugees and Others of Concern to UNHCR, 1999 Statistical Overview and OECD Trends in International Migration

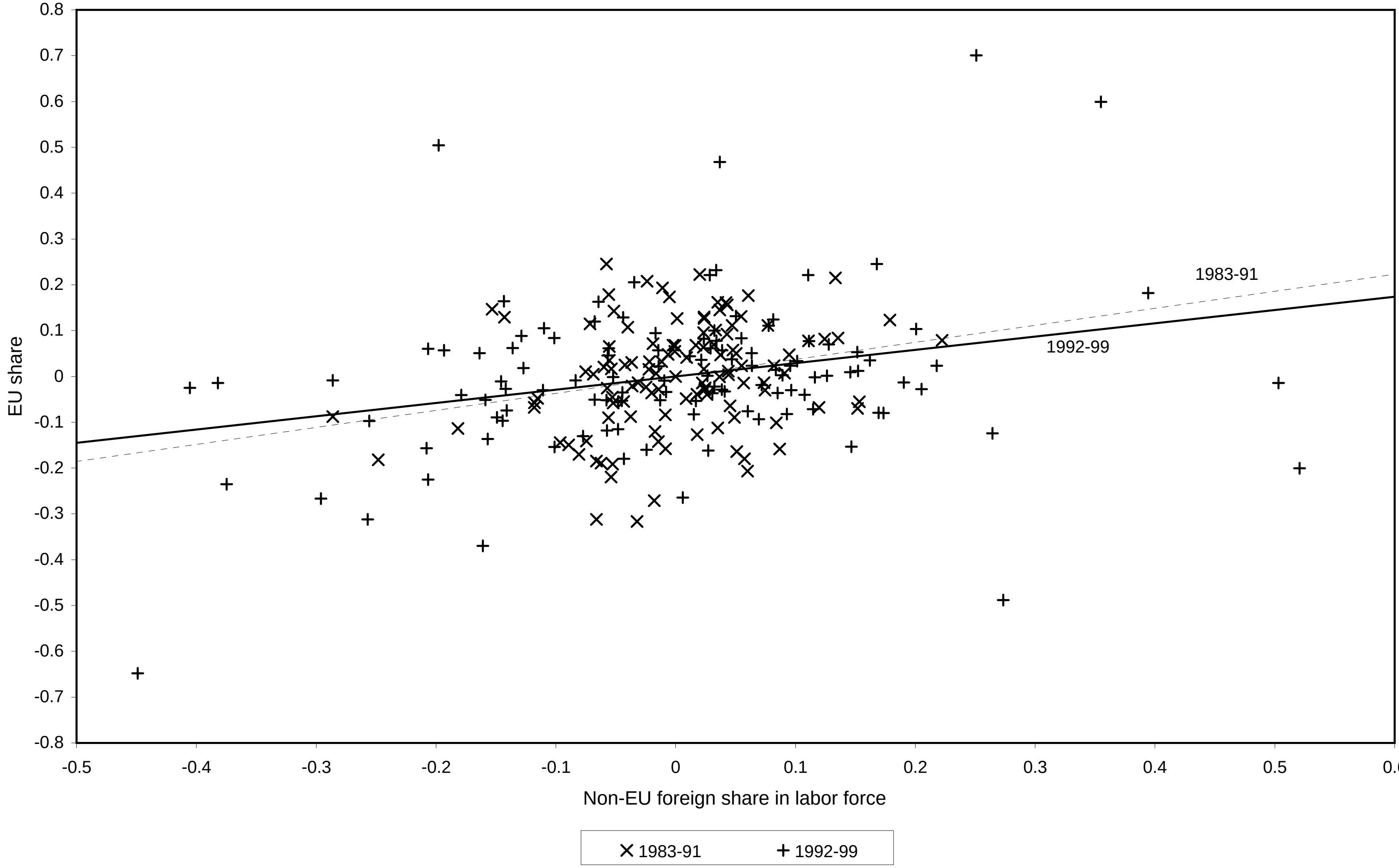

Figure 3. EU foreign share plotted against non-EU foreign share,
adjusted for country and year effects.

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