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PROTECTIVE OR COUNTER-PRODUCTIVE? LABOUR MARKET INSTITUTIONS AND THE EFFECT OF IMMIGRATION ON EU NATIVES*

Joshua D. Angrist and Adriana D. Kugler

Reduced labour market flexibility may protect some native workers from immigrant competition but can increase negative effects on equilibrium employment. This motivates an analysis of immigration effects interacted with institutions. OLS estimates for European countries show small, mostly negative immigration effects while an IV strategy based on immigrants from former Yugoslavia generates larger though mostly insignificant negative estimates. Specifications allowing interactions between immigration and measures of labour and product market rigidity are consistent with the view that reduced flexibility increases negative immigration effects. The estimates typically imply more native job losses in countries with restrictive institutions, especially restricted product markets.

The European immigration policy debate is fuelled by the fact that immigration now accounts for the bulk of population growth in the European Union (EU) (OECD, 1999*a*). Many observers also note 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 has been associated with high levels of anti-foreigner sentiment, and the view that immigrants take jobs from natives is widespread (Bauer *et al.*, 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 (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 been modest.[1]

This paper takes a fresh look at the employment consequences of immigration in Western Europe, motivated by two considerations. First, we use the Balkan Wars as a Mariel-Boatlift style immigration experiment along the lines of Card's (1990). Second, we focus on institutional aspects of the immigration question. Many observers have argued that persistent high unemployment in Europe is due to institutions that increase turnover and employment costs, e.g. OECD (1994).

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[1] European studies include Pischke and Velling (1997) for Germany; Winter-Ebmer and Zweimuller (1997) for Austria, and Dolado *et al.* (1996) for Spain. Similar to Card' (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.

Recently, Blanchard and Wolfers (2000) extended this inquiry by suggesting that the negative employment consequences of a rigid labour market are felt not so much in good times, but rather in the labour market's response to adverse demand shocks. Here, we integrate the institutions debate with an analysis of shocks that occur on the labour-supply side. In particular, we ask whether institutional characteristics, such as employment protection, high replacement rates, rigid wages, and high business entry costs, affect the employment consequences of immigration-induced increases in the work force.

While economic 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 labour and product markets. In fact, Rodrik (1997) has argued that the demand for social insurance is in part a response to the pressures of global economic integration, including increased migration. However, the equilibrium consequences of protective regulations and institutions are unclear. We therefore begin with a theoretical discussion of immigration interactions with institutions. Although employment protection and entry barriers may reduce job loss in the short run, our theoretical discussion shows why reduced flexibility may be counter-productive, possibly amplifying any negative employment consequences of immigration for natives.[2]

The empirical relationship of interest in this study is the effect of immigrant shares on native employment rates, where this effect is allowed to vary with institutional characteristics. The estimates use a panel data set for 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 labour market variables by age, sex, education, and nationality or nativity. This data set allows us to conduct analyses much like previous immigration research for the US and individual European countries using micro data, while allowing consistent cross-country comparisons. In addition to exploring interactions with institutions, our cross-country analysis may address some of the methodological questions raised in the previous literature on immigration effects. First, cross-country data may be less affected by endogenous mobility than data on US states or cities. Second, bias from endogenous mobility may be mitigated by use of the two 1990s' Balkan Wars (in Bosnia and Kosovo) as a source of exogenous variation providing shocks to immigrant flows in Europe.

1. Theoretical Framework

A competitive model with two types of labour and exogenous separations illustrates standard predictions regarding the impact of immigrants on natives, and shows how the effect of immigrants on native employment might be modified by

[2] Previous theoretical studies of interactions between immigration effects and labour market flexibility include Schmidt *et al.* (1994) and Razin and Sadka (1996). Kugler and Saint-Paul (2003) look at the interaction between employment protection and the re-employment prospects of workers subject to individual shocks.

[3] Our sample is not exactly the EEA. It includes the EU 15 plus Iceland and Norway, omits Liechtenstein but includes Switzerland, which opted out of the EEA.

institutions such as firing costs. Although our theoretical framework does not lead to a detailed structural specification, it serves to motivate the empirical work. The setup is similar to that appearing in earlier analyses of immigration questions, augmented with elements used by Acemoglu and Angrist (2001) and Saint-Paul (1996) to study the effects of labour market regulation and employment protection. The main predictions are that immigration tends to reduce native employment if natives and immigrants are at least moderately substitutable in production; labour market institutions that reduce the employment of natives but not immigrants may exacerbate the short-run negative impact of immigration; and finally that barriers to firm entry slow the rate at which employment returns to pre-immigration levels.

To make these points, firm output is assumed to be produced by immigrants and natives with production function,

$$f[\theta g(N_t, I_t)];$$

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

where N_t is the number of natives (or nationals) and I_t is the number of immigrants (or non-nationals). The function $g(N_t, I_t)$ is a CES-type labour aggregate as in Card (2001) and θ is a location-specific shifter as in Lalonde and Topel (1991). The first derivative of the production function is positive and the second derivative is negative, reflecting the presence of inputs other than the labour aggregate. Our approach differs modestly from others in the literature in that we specify CES interaction between immigrants and natives as a group. In contrast, immigrant-native complementarity in Altonji and Card (1991) is generated by differences in skill or education, with groups at the same skill level being perfect substitutes. Language differences make this harder to motivate in the European context, while at the same time immigrants to the EU are not dramatically less educated than natives.

An important feature of many European labour markets is high firing costs. These come in the form of bureaucratic limitations on dismissals, requirements for severance pay, and restrictive collective bargaining agreements. Immigrants are probably less likely than natives to be covered by these provisions, however, 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 labour force becomes unproductive in the current job, and is therefore laid off.[4] Immigrants and natives are paid real wages, w_{Nt} and w_{It}, with the price of output as numeraire.

[4] As in Acemoglu and Angrist (2001), productivity declines are assumed to be large 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).

Immigration effects are derived in a simple dynamic setup where price-taking firms act to maximise the present value of profits, with discount factor φ. In particular, firms' objective functions can be written,

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

Adjustment costs are linear and there is no aggregate uncertainty, so time subscripts can be dropped and the objective simplified to be:

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

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 (2*a*), which implicitly defines the demand function for native labour, 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, c_N, are now defined as a proportional to the native wage, in line with the specification of firing costs in many countries. In Spain, for example, unjust dismissal costs are set at about 12% of annual earnings.

The labour 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 labour 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 labour supply seems reasonable since natives are more likely than immigrants to have access to social insurance. The native labour supply function is

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

where P is the native population, r is the unemployment insurance replacement rate, and ε is the native labour supply elasticity, assumed to be positive. For what follows, it is useful to define the inverse labour 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 obtained by allowing m to be endogenous and determined by the condition that profits are equal to entry costs. In the short run, (2*a*), (2*b*), and (3) determine the two endogenous wage levels and the number of employed natives. Since immigrant labour supply is exogenous, the key equilibrium condition can be written as follows

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

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

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

Equation (4) provides the basis for our empirical work, which relates the log of native employment to the log of the immigrant share in the labour force. Following Lalonde and Topel (1991), we think of the estimates as approximating (4) to first order in logs. The estimating equation is assumed to hold at the country level since firms are taken to be identical except for an additive random error and factors that can be absorbed by country and year effects. Before turning to the empirical results, the theoretical framework is used to highlight possible interactions between institutional characteristics and immigration effects. We start with 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 then consider long-run impacts and the interaction between immigration and product market restrictions.

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

$$\mathrm{dln}\, N^*/\mathrm{dln}\, M = (\partial N/\partial I)(I/N) = \xi_{NI}(\varepsilon^{-1} - \xi_{NN})^{-1} \equiv e(N, \varepsilon), \tag{5}$$

where ξ_{NI} and ξ_{NN} are the elasticities of factor price for native wage rates with respect to immigrant and native employment. That is, $\xi_{NI} = (\partial w_N/\partial I)(I/w_N)$ and $\xi_{NN} = (\partial w_N/\partial N)(N/w_N)$, as determined by the demand curve for native labour. The notation $e(N, \varepsilon)$ is used as shorthand for $\mathrm{dln}N^*/\mathrm{dln}M$ to emphasise the fact that parameters other than the labour supply elasticity modify the immigrant impact only through native employment levels. This expression, derived in our working paper (Angrist and Kugler, 2001), is similar to the corresponding relation in Johnson's (1980) static model, though here immigration has an ambiguous effect on native employment. The own-wage term ξ_{NN} is negative, so the denominator is positive, but ξ_{NI} in the numerator can be positive or negative depending on the extent of immigrant-native complementarity.

Immigration is predicted to reduce native wage rates for most plausible parameter values in this sort of model, so that $e(N, \varepsilon)$ is negative (Altonji and Card, 1991; Ichino, 1993). Of course, if immigrants and natives are perfect substitutes ($\rho = 1$), then ξ_{NI} and $e(N, \varepsilon)$ are necessarily negative. More generally, immigrants displace natives as long as the elasticity of substitution between immigrants and natives, $1/(1 - \rho)$, is above the Cobb–Douglass benchmark of 1 (or even less than 1 if demand for natives is less than unit elastic). Assuming technological parameters are in this range, immigration reduces native employment, with larger effects when native labour supply is more elastic.

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

$$\partial e/\partial c_N = (\partial e/\partial N)(\partial N/\partial c_N).$$

It is clear from the equilibrium condition, (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).[5] 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, expand $\partial e/\partial N$ 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 is likely to 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, aggregate labour demand becomes less elastic as the labour share increases, so ξ_{NN} becomes more negative as N grows (Hamermesh, 1986).

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

$$\partial e/\partial r = (\partial e/\partial N)(\partial N/\partial r).$$

As with firing costs, higher replacement rates reduce native employment levels. In particular,

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

Higher replacement rates therefore increase any job loss due to immigration if scale effects are positive. 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.

In practice, scale effects are likely to be hard to detect unless the overall effect of restrictive institutions is substantial. Suppose, for example, as suggested by Lazear's (1990) estimates, that employment protection at the level found in Spain reduces employment levels by 6%. If scale effects are such that a 10% reduction in employment increases the job loss from immigrant competition by 20% (and this seems like a large effect), then employment protection at the Spanish level increases immigration-related job losses by 12%. Although not trivial, this is small relative to the precision with which we typically expect to be able to measure immigration effects.

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

A second and probably more important channel for institutional interactions in the European context is wage rigidity. We omit a detailed analysis of the impact of union wage setting or minimum wages, but look briefly at a stylised characterisation 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 potentially a large effect. For example, if $\xi_{NN} = -1$ and $\varepsilon = \frac{1}{2}$ in the absence of restrictions, immigration-induced job losses are three times greater with rigid wages. Moreover a higher w_N reduces employment even without immigration, leading to the same sort of scale effect discussed earlier for firing costs and replacement rates. We should note, however that our brief analysis of wage rigidity omits any off-setting feedback effects whereby union wage demands are moderated as a consequence of competition from immigrants, a possibility discussed by Schmidt *et al.* (1994).

1.1. *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 analyse the effect of immigration on profits. Using the envelope theorem, 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. 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 paid his or her value of marginal product, infra-marginal workers are paid less.[6]

Assuming profits were equal to entry costs before immigration and there are diminishing returns to labour 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. More formally, the effect of immigration on aggregate employment is shown in our working paper to be

[6] The exception is if there are constant returns for labour inputs alone, i.e., $f'(\cdot) = 0$, in which case profits are always zero and there is no entry. Even if native wages rise due to immigrant-native complementarity, profits increase as long as the production function exhibits diminishing returns.

$$(\partial N^*/\partial M)(M/N^*) = e(N,\varepsilon)(1 - \partial \ln m/\partial \ln M) - \xi_{NN}(\varepsilon^{-1} - \xi_{NN})^{-1}(\partial \ln m/\partial \ln M), \tag{6}$$

where $e(N, \varepsilon)$ is the short-run employment response defined in (5). Since $\xi_{NN} < 0$ and $\partial \ln m/\partial \ln M \leq 1$, 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 immigration on native employment must be 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 further aggravate job losses from immigration. This is because factors that increase labour costs will also tend to reduce or slow new firms' entry in response to low-cost immigrant labour. Finally, entry costs that reduce native employment levels will interact negatively with other rigidities because of the short-run scale effect noted above.[7]

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 labour market flexibility, replacement rates and barriers to entrepreneurship. To establish a baseline, however, the empirical discussion begins with a reduced-form analysis of immigration effects that omits interactions with institutional characteristics.

2. Background and Data

2.1. *Descriptive Statistics*

In recent years, the European countries with the largest proportion of labour force from non-EU countries have been Austria, France, Germany, Sweden, Switzerland, and the UK. This can be seen in Table 1, which reports descriptive statistics from the Eurostat labour force surveys for 18 EU and other EEA countries.[8] 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 ethnically similar migrants. Germany, for example, accepts large

[7] See Bertrand and Kramarz (2001) for recent evidence on the employment consequences of entry costs in the retail industry in France. Layard and Nickell (1999) emphasise the interaction between restricted product market entry and union bargaining power.

[8] Additional information on data sources and extracts is provided in the Data Appendix.

Table 1

Descriptive Statistics

Country	LFS Coverage	Labour force in 1999 (000s)	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
UK	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. All statistics are for men and women aged 20–59 in the Eurostat Labour Force Survey.

numbers of ethnic Germans, mostly from the former Soviet Union. Similarly, roughly 45% of the foreign population in Greece in 1997 was of Greek background, mostly from the former Soviet Union and Albania (OECD, 1999*a*).

Figure 1 documents the time pattern of immigration for many of the important immigrant-receiving countries. There was little change in immigration up to 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. 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 in the Kosovo War. Figure 2 also shows that Yugoslavs accounted for more than 30% of asylum-seekers in the war years. Yugoslav asylum seekers were a significant part of the total foreign inflow, with wartime modes at 10–15%. Since many foreigners in EU countries come from other EU countries, the effect of the Yugoslav asylum seekers on the non-EU foreign

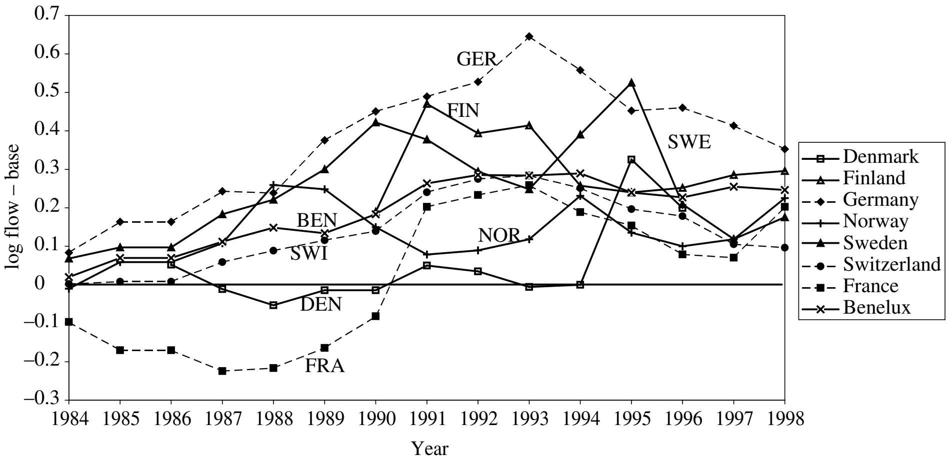

Fig. 1. *Foreign Inflows 1984–98.* The figure shows log counts minus a 1983 base. Data are from population registers except for France. Data sources are given in the Appendix.

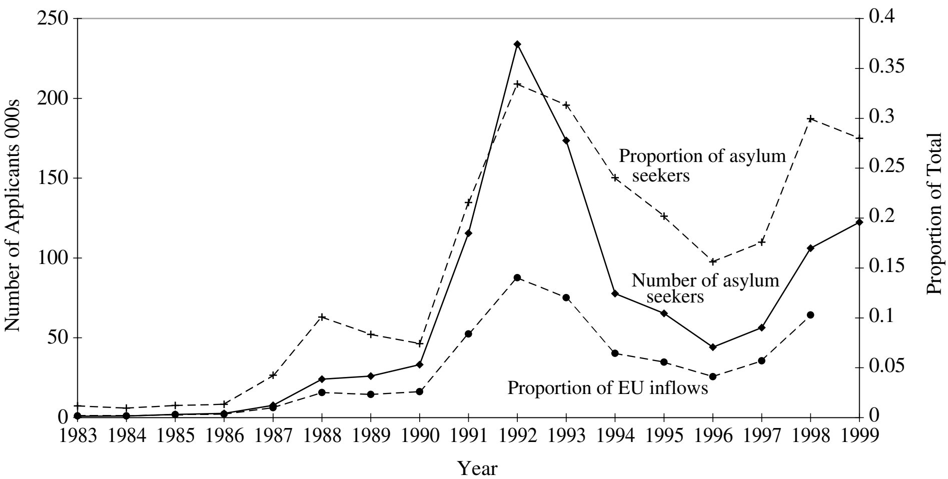

Fig. 2. *Number of Yugoslav asylum applicants in Europe,* Yugoslav proportion of total applicants, and Yugoslavs as a proportion of total foreign inflows to selected EU countries. Data for 1992–99 include asylum applicants from both Yugoslavia and Bosnia. Data sources are given in the Appendix.

share is considerably larger than indicated by the Figure. Our data suggest that in the 1995–9 period, roughly one-third of male immigrants aged 20–59 from non-EU countries were from former Yugoslavia.[9]

The importance of migration within the EU is also documented in Table 1. In some countries, many non-national residents in the labour force are from other developed European nations. Luxembourg is a clear outlier, with over one-third of

[9] See also Table I.2 in OECD (2001), which shows Yugoslavia to be among the top 5 immigrant source countries for the immigrant stock in a number of OECD states in 1998.

its labour force from other EU countries. France is more typical, with 2–3% of its labour 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.

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'. We briefly explored the overlap between nationality and nativity-based definitions of immigrant status in 1996, the middle of the period for which data on country of birth are available (1992–9). Almost all of those with non-EU nationality were foreign born, so that at least among adults, most non-nationals are indeed immigrants.[10] Perhaps surprisingly, however, naturalisation rates for non-Europeans are substantial. In France, for example, about half of residents aged 20-59 who were born in a non-EU country had obtained French citizenship. On the other hand, few *recent arrivals* were naturalised, consistent with the fact that many EU countries impose an extensive waiting period before naturalisation. Although there are a few anomalies (e.g, the high naturalisation rates for ethnic Greeks and ethnic Germans), the group of non-nationals can be seen as roughly coincident with the group of recently-arrived foreign born residents. Recently arrived immigrants are of special interest here since the instrumental variables strategy primarily captures the effects of recent migrants from the Balkans.[11]

2.2. *Immigrant Employment Policies*

Before turning to the estimates, we summarise government policies and statistics relating to the ease of labour-market access and legal employment opportunities for immigrants to EU countries. The OECD migration volumes describe aspects of immigration policies in the EU, for example, OECD (1999*a*). Of special relevance for our study is the treatment of immigrants and asylum-seekers from Yugoslavia. 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 had 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 in Germany, 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 fuelled by immigrants, in spite of the fact that an estimated 89% of asylum seekers entered Italy illegally in recent years.

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

[11] Also relevant is the possibility, noted by Card (2001), that recent immigrants compete more directly with low-skilled natives than veteran immigrants.

In Sweden in 1998, the largest immigrant group in the labour force after the Finns were those from former Yugoslavia. A special visa programme for Bosnians and other parts of the former Yugoslavia operated in Sweden in 1993–6. 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 labour force. This can be seen in Table 2, which reports labour-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 1995–9 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 labour force in every country. The contrast between natives and recent immigrants 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% of veteran Yugoslav men and 81% of other veteran non-EU men were in the labour force, while 65% of recent Yugoslav men and 62% of other recent non-EU men were in the labour 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 2, which shows 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% for men in the UK, to rates above 75% 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 over time.

Also of interest is the relative education levels of immigrants and natives, since there is probably more competition for jobs within schooling groups than between groups. The LFS extract includes information for some years and countries on the size of three schooling groups. Although immigrants are less educated than natives, the education statistics show considerable overlap between the immigrant and native schooling distributions, at least for men. For example, about a third of native men had a 9th grade education or less, while 42% of non-EU immigrant men were in this category. For details, see Angrist and Kugler (2001).

Table 2

Characteristics of Immigrants and Natives

	Men					Women				
		Non-EU		Yugoslavs			Non-EU		Yugoslavs	
Country	Natives (1)	Recent (2)	Veteran (3)	Recent (4)	Veteran (5)	Natives (6)	Recent (7)	Veteran (8)	Recent (9)	Veteran (10)
A. Labour 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
UK	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
UK	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

Table 2

Continued

	Men					Women				
		Non-EU		Yugoslavs			Non-EU		Yugoslavs	
Country	Natives (1)	Recent (2)	Veteran (3)	Recent (4)	Veteran (5)	Natives (6)	Recent (7)	Veteran (8)	Recent (9)	Veteran (10)
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. Data are unavailable for Switzerland and unreliable for Iceland. A dash denotes entries below the LFS confidentiality threshold.

3. Estimates of Immigration Effects

3.1. *OLS Estimates*

We begin with models allowing a single immigration effect for all countries, turning afterwards to models that incorporate interactions with institutions. The first equation estimated is

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

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 the immigrant (non-national) proportion of the labour force. Measuring the immigrant share by proportion of the labour force instead of proportion of population partly adjusts for composition effects not fully captured by demographic controls and for the fact that employment rates for immigrants may be understated.

Equation (7) can be interpreted as approximating the first-order condition determining native employment or 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 labour demand shocks correlated with both immigrant shares and native employment. We therefore experimented with models that include controls for the log of the foreign share with EU nationality, denoted $\ln(u_{jt})$. This provides a partial though potentially endogenous control for local demand factors that may increase overall immigration. Including $\ln(u_{jt})$ also addresses the point, raised in earlier immigration studies, that internal migration flows act to offset the impact of immigrants; see Card (2001) for a discussion of this possibility in the US context. Internal EU migrants are probably similar in many respects to US internal migrants, in that they are drawn to host countries by job assignments and employment opportunities. Moreover, the fact that $\ln(u_{jt})$ is a potentially endogenous control should not bias the IV estimates since the instruments are uncorrelated with migration from other EU countries.

OLS estimates of models for native men that omit the EU share, reported in Panel *A* of Table 3, show no significant effect of non-EU immigrants overall, though there is a small and significant effect of −0.021 when the sample is limited to young natives. Including the EU share as a control variable leads to an even larger negative effect of −0.037 for young native men, and a significant effect of −0.021 overall. These results are reported in Panel *B* of Table 3, cols. 1–3. The larger estimates generated by models that include the EU share are due to the fact that the EU and non-EU foreign shares are positively correlated. In practice, therefore, while migration within the EU is positively correlated with native employment, internal EU migration does not act to 'undo' the effects of non-EU migrants. Taking the pooled significant estimate from the model with EU share as representative of the impact on men, the magnitudes are such that 100 immigrants

Table 3
OLS Estimates

		No trends			With trends		
			By age group			By age group	
Non-national share	Sex	Pooled (1)	Under 40 (2)	Over 40 (3)	Pooled (4)	Under 40 (5)	Over 40 (6)
A. Without share of Non-Nationals from EU							
Non-EU	Men	−0.010 (0.007)	−0.021 (0.007)	0.002 (0.004)	−0.009 (0.013)	−0.011 (0.013)	−0.007 (0.006)
Non-EU	Women	0.0002 (0.021)	0.002 (0.013)	−0.001 (0.022)	−0.012 (0.034)	−0.022 (0.013)	−0.003 (0.012)
N		420	210	210	420	210	210
B. With share of Non-Nationals from EU							
Non-EU	Men	−0.021 (0.008)	−0.037 (0.008)	−0.004 (0.005)	−0.011 (0.015)	−0.012 (0.012)	−0.010 (0.007)
EU		0.036 (0.016)	0.053 (0.014)	0.018 (0.010)	0.022 (0.019)	0.028 (0.009)	0.016 (0.006)
Non-EU	Women	−0.026 (0.026)	−0.026 (0.016)	−0.026 (0.026)	−0.012 (0.048)	−0.023 (0.012)	−0.002 (0.015)
EU		0.086 (0.041)	0.092 (0.026)	0.081 (0.031)	0.008 (0.049)	0.018 (0.011)	−0.002 (0.013)
N		402	201	201	402	201	201

Notes: The Table reports OLS estimates of (7) in the text. Robust standard errors are reported in parentheses. Estimates in cols. 4–6 are from models that include country specific trends.

in the labour force cost about 35 native jobs in a country where 5% of the labour force is non-EU foreign (as Table 1 indicates for Germany in 1989). In contrast with the results for men, none of the estimates for women in cols. 1–3 of Table 3 are significant.

Equation (7) relies on time-invariant country effects and/or $\ln(u_{jt})$ to control for omitted variables correlated with immigration rates. OLS estimates of the parameters in this equation are biased if immigration is correlated with country-specific trends, a problem made more likely by the long time-series sample. We therefore also report results in cols. 4–6 from models replacing the country effect, β_j, with a country-specific linear trend, $\beta_{0j} + \beta_{1j}t$. It should be noted, however, that robustness to the inclusion of country-specific trends is a stringent test, since any local trend or near-trend component in immigration is removed in this specification.

The negative employment effects for younger women become slightly larger and borderline 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. Of course, the inclusion of country trends does not necessarily eliminate bias from endogenous migration and it makes any immigration effects harder to detect since the resulting estimates are typically less precise.

3.2. *IV Estimates*

The OLS estimates in Table 3 may be biased upwards by immigrants choosing to locate where their employment prospects are best. This Section discusses estimates of immigration effects 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.

The first-stage equation for the IV estimates is

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

where

b_{jt} = distance from Sarajevo $\times$ dummy for 1991–5 (Bosnia War years)

n_{jt} = distance from Sarajevo $\times$ dummy for 1996–7 (inter-war years)

k_{jt} = distance from Pristina $\times$ dummy for 1998–9 (Kosovo War years)

are the excluded instruments, and τ_t and ψ_j are year and country effects. The distance from potential host countries is measured as approximate miles either from the nearest city with a population of at least 100,000 or from the capital.

The essence of the IV strategy is to look for a break in the time-series behaviour of employment rates for countries relatively close to Yugoslavia. Therefore, as a specification check, we also estimated IV models with a parametric control for country-specific linear trends, as with the OLS estimates. 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. As before, some specifications include the EU foreign share, i.e. $\ln(u_{jt})$, as a covariate. As we show below, however, this has little effect on the IV estimates because the EU share is largely uncorrelated with the instruments.

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 4, which reports the coefficients on b_{jt}, n_{jt}, and k_{jt}, plus a pre-war interaction term as a specification check. The coefficients are scaled so that they represent the effect of 1,000 miles. For example, 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% (see cols. 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 dummies 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* show 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 4 reports the results of replacing the non-EU foreign share with the EU foreign share as the dependent variable in (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* indeed reflect the effect of immigrants from former Yugoslavia. Moreover, it means that the IV estimation strategy is unaffected by inclusion of 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 large cities. Because the first-stage relationship is stronger when distance is measured from large cities, we used this variable to construct the second-stage estimates.

The 2SLS estimates using b_{jt}, n_{jt}, and k_{jt} as instruments are reported in Table 5, separately for models that do and do not control for $\ln(u_{jt})$. For men, the estimated immigration effects without the EU share are on the order of −0.05 when age groups are pooled and −0.08 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, some of 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 −0.05 as

Table 4

First Stage Estimates

	Distance to nearest big city				Distance to capital			
	No trends		With trends		No trends		With trends	
Instruments	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Non-Nationals from Non-EU countries								
Bosnia War (1991–5) × 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)
Inter War (1996–7) × 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 (1998–9) × 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)
Pre-war (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)
Inter 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)
Pre-war × 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

Notes: The Table reports the effect of the Balkan War periods interacted with the distance from Sarajevo or Pristina on the log share of non-nationals from non-EU and EU countries. All models include country and year effects. The sample size is 201 for non-nationals from non-EU countries and 202 for non-nationals from EU countries.

Table 5
IV Estimates

		No trends			With trends		
			By age group			By age group	
Non-national share	Sex	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 (0.023)	−0.082 (0.030)	−0.018 (0.016)	−0.034 (0.029)	−0.028 (0.027)	−0.040 (0.018)
Non-EU	Women	−0.245 (0.093)	−0.189 (0.070)	−0.301 (0.102)	−0.058 (0.112)	−0.034 (0.030)	−0.082 (0.046)
N		420	210	210	420	210	210
B. With share of Non-Nationals from EU countries							
Non-EU	Men	−0.050 (0.016)	−0.089 (0.020)	−0.011 (0.011)	−0.043 (0.034)	−0.042 (0.031)	−0.045 (0.020)
EU		0.059 (0.016)	0.094 (0.020)	0.024 (0.012)	0.032 (0.017)	0.037 (0.016)	0.027 (0.010)
Non-EU	Women	−0.210 (0.064)	−0.166 (0.043)	−0.253 (0.062)	−0.067 (0.132)	−0.046 (0.033)	−0.088 (0.053)
EU		0.232 (0.066)	0.203 (0.044)	0.260 (0.064)	0.026 (0.067)	0.025 (0.017)	0.026 (0.027)
N		402	201	201	402	201	201

Notes: The Table reports IV estimates of (7) in the text, with the non-EU share endogenous. The EU share variable in panel *B* is treated as exogenous.

representative, the IV estimates for men predict that increasing the foreign share by 10% would reduce employment by half of a percent in a country where 5% of the labour force is foreign. On a per-worker basis, this implies that 100 immigrants in the labour 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. Hunt (1992) also finds large effects in her study of repatriates from Algeria to France, a country with very restrictive labour and product market regulations. It should be noted, however, that the estimates with country trends are mostly smaller than those without trends and not significant.

While many of the 2SLS estimates for men are imprecise, they suggest a pattern of reasonably stable negative 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 larger still. Note, however, that a given percentage effect for women translates into a smaller effect on levels than would do for men because of lower female labour 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 labour 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.[12]

4. Immigrants Interact with Institutions

Do institutions that make labour and product markets more rigid or less competitive change the employment consequences of immigration for natives?

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

The theoretical section suggests that restrictive labour 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 may worsen the employment impact of immigrants because of scale effects and especially because rigid wages make native workers less competitive with immigrants. 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 neutralise any displacement effects. Finally, entry costs that reduce employment levels also have an adverse 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 variables correlated with the instruments. We therefore continue to present results that control for linear country trends, as well as OLS estimates. It seems reasonable to think of the OLS and IV estimates as roughly bounding the effects of interest and to look for a consistent pattern of interaction terms in the two sets.

Our empirical strategy looks at OLS and IV estimates of interactions with measures of three of the institutional features discussed in the theory section. The first is a summary of labour standards that indexes the extent of employment protection, restrictions on work hours and employment contracts, administrative or union oversight in hiring and firing decisions, and minimum wages. This measure therefore reflects both firing costs and wage rigidity. The second is the average replacement rate. Both variables were taken from Table 4 in Nickell (1997). Labour standards are captured by an index ranging from 0–7, with 7 denoting the most restrictive institutions in our sample. 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 *et al.* (2000). This is an index of barriers to entrepreneurship ranging from 0.5 to 2.75 in our sample.

4.1. *Estimates of Interaction Effects*

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

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

where $\tilde{x}_j$ is an institutional characteristic, measured as the deviation from the median value in our sample. We also include the EU share since this increased the precision of the IV estimates in Table 5 in models without country trends. To increase the comparability of estimated effects, institutional variables are scaled in standard deviation units. 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 with a one standard deviation change in $\tilde{x}_j$. We think of α_{1i} as approximating the average derivative of $e(N, \varepsilon)$ with respect to institutional variables. When $\ln(s_{jt})$ is treated as endogenous, the instrument list is augmented with interactions between $\tilde{x}_j$ and the instruments used to estimate (7). Note that this setup fails to identify the effects of the institutions themselves since $\tilde{x}_j$ is time-invariant.[13]

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 and the identification of interaction terms in models with country trends is tenuous. As noted earlier, this sensitivity appears to be due to strong regional trends in female labour force participation. These trends vary across countries in a manner correlated with distance from former Yugoslavia.

The estimates of (9) are consistent with the hypothesis that immigration effects are more negative in countries with less flexible labour markets, higher replacement rates, and higher entry costs. This can be seen in Panel *A* of Table 6, which labels estimates of α_{0i} 'Main Effect' and estimates of α_{1i} 'Interaction'. The first column shows OLS results for men in both age groups. The interaction with labour standards in this specification is estimated to be a statistically significant −0.015, indicating that increasing the severity of labour standards by one standard deviation, about the difference between Denmark and Belgium, would increase the negative effect of immigration from −0.027 at the median to about −0.042. The interaction terms are larger for young men than for men over 40. Similarly, the pooled interaction with replacement rates is −0.017, so a one standard deviation increase in replacement rates would increase the negative effect of immigration from −0.027 at the median to −0.044. The results including both labour 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 younger men.

The 2SLS estimates for men, reported in cols. 4–6, 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 labour standards. The 2SLS estimates of the interaction terms for labour standards 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. The 2SLS estimates of models incorporating interactions with both

[13] Direct causal effects of labour market institutions are difficult to capture in a cross-country panel because of a lack of variation. An OECD (1999*b*, p. 50) report observes, 'Between the 1980s and late 1990s, there was considerable continuity in EPL practices in most countries.' There have been a few policy experiments, however. Kugler *et al.* (2002) find increased employment in response to a recent sharp decrease in firing costs and payroll taxes for some protected groups in Spain. For a dissenting view on the importance of employment protection, see Layard and Nickell (1999), who draw conclusions from a cross-sectional analysis.

Table 6

Interactions with Institutions: Estimates for Men

Interaction	Regressor	OLS			2SLS		
			By age group			By age group	
		Pooled (1)	Under 40 (2)	Over 40 (3)	Pooled (4)	Under 40 (5)	Over 40 (6)
A. Institutions one at a time							
Labour standards	Main effect	−0.027 (0.010)	−0.045 (0.012)	−0.010 (0.008)	−0.031 (0.025)	−0.071 (0.034)	0.008 (0.021)
	Interaction	−0.015 (0.007)	−0.019 (0.008)	−0.011 (0.005)	−0.070 (0.025)	−0.091 (0.034)	−0.050 (0.021)
N		334	167	167	334	167	167
Replacement rate	Main effect	−0.027 (0.010)	−0.045 (0.012)	−0.010 (0.007)	0.051 (0.056)	0.108 (0.089)	−0.006 (0.040)
	Interaction	−0.017 (0.008)	−0.019 (0.009)	−0.014 (0.006)	0.007 (0.017)	0.0001 (0.027)	0.015 (0.012)
N		334	167	167	334	167	167
Barriers to entrepreneurship	Main effect	−0.027 (0.008)	−0.044 (0.009)	−0.010 (0.006)	−0.061 (0.020)	−0.117 (0.028)	−0.005 (0.017)
	Interaction	−0.019 (0.009)	−0.023 (0.011)	−0.014 (0.007)	−0.039 (0.022)	−0.020 (0.031)	−0.058 (0.019)
N		368	184	184	368	184	184
B. Institutions Together							
Labour standards and replacement rate	Main effect	−0.028 (0.010)	−0.046 (0.012)	−0.011 (0.007)	−0.032 (0.029)	−0.080 (0.042)	0.016 (0.021)
	Labour standards	−0.012 (0.007)	−0.016 (0.008)	−0.008 (0.005)	−0.056 (0.020)	−0.091 (0.030)	−0.020 (0.015)
	Replacement rate	−0.013 (0.008)	−0.015 (0.009)	−0.012 (0.006)	−0.016 (0.017)	−0.029 (0.026)	−0.002 (0.013)
N		334	167	167	334	167	167
Barriers, labour standards, and replacement rate	Main effect	−0.035 (0.012)	−0.053 (0.014)	−0.017 (0.008)	−0.055 (0.033)	−0.102 (0.044)	−0.008 (0.026)
	Barriers	−0.017 (0.015)	−0.018 (0.018)	−0.016 (0.011)	−0.088 (0.032)	−0.124 (0.043)	−0.052 (0.025)
	Labour standards	−0.002 (0.011)	−0.006 (0.013)	0.001 (0.008)	0.024 (0.020)	0.029 (0.026)	0.020 (0.015)
	Replacement rate	−0.016 (0.008)	−0.017 (0.010)	−0.014 (0.006)	0.006 (0.013)	0.006 (0.018)	0.007 (0.010)
N		334	167	167	334	167	167

Notes: The Table reports main effects and interaction terms in (9) in the text. Instruments for the foreign share are as in Table 5, plus interactions with institutional measures. Main effects are evaluated at the median institution (5 for labour standards, 63 for replacement rate, 1.715 for barriers to entrepreneurship) and measured in standard deviation units. The EU share variable is included and treated as exogenous.

labour standards and replacement rates still show negative and mostly significant interaction terms for labour standards, with insignificant negative interactions for replacement rates.

The stronger evidence of negative interactions with labour standards than with replacement rates is consistent with our theoretical story, which attributes interactions with replacement rates solely to scale effects. In principle, the labour standards variable also captures an element of wage rigidity, which is theoretically likely to interact more strongly with immigration to reduce employment.

As with the measures of labour market flexibility, the results of estimating models allowing interactions with barriers to entry show immigration effects that are more negative in countries with higher barriers. For example, the pooled estimate in col. (1) suggests that increasing entry barriers by one standard deviation, about the difference between Germany and France, would increase the negative effect of immigration from −0.027 at the median to −0.046. The interactions with entry costs are again larger for men under 40, and larger when estimated by 2SLS.

The results of the attempt to estimate the interaction with entry costs jointly with interactions with labour market flexibility generates results less clear cut than when estimated individually. OLS estimates of interactions with entry costs are similar to those estimated one at a time, though no longer significant. Interactions with replacement rates also remain significant and negative in this specification, though they are smaller than the estimates of interactions with labour standards. The corresponding 2SLS estimates of interactions with entry barriers in models with multiple characteristics are much larger (i.e., more negative), perhaps implausibly so, though the interaction-term standard errors are also large. In fact, the interactions with entry barriers are the only 2SLS estimates significant in models with multiple characteristics.

While our ability to distinguish specific institutional mechanisms is limited, the results in Table 6 show a pattern of negative interactions between immigration and institutional variables that reflect reduced labour and product market flexibility. As a final check on these estimates, we added country trends, as in the models without institutional interactions. This necessarily leads to a substantial loss of precision since country-specific trends are also interaction terms.

The results with country trends, reported in Table 7, nevertheless show mostly negative and sometimes significant interactions, with 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 labour 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. On balance, the interaction with barriers to entry is perhaps the most robust finding arising from the institutional analysis. The relative importance of interactions with labour standards and replacement rates varies from specification to specification but labour standards are consistently more important in models that omit country trends.

Table 7

Interactions with Institutions: Estimates for Men with Country Trends

		OLS			2SLS		
			By age group			By age group	
Interaction	Regressor	Pooled (1)	Under 40 (2)	Over 40 (3)	Pooled (4)	Under 40 (5)	Over 40 (6)
A. Institutions one at a time							
Labor standards	Main effect	−0.014 (0.016)	−0.019 (0.014)	−0.010 (0.009)	−0.050 (0.064)	−0.067 (0.064)	−0.032 (0.038)
	Interaction	−0.003 (0.011)	−0.006 (0.010)	−0.0002 (0.006)	−0.034 (0.037)	−0.051 (0.037)	−0.018 (0.022)
N		334	167	167	334	167	167
Replacement rate	Main effect	−0.022 (0.016)	−0.027 (0.014)	−0.018 (0.009)	−0.045 (0.066)	−0.003 (0.062)	−0.086 (0.044)
	Interaction	−0.026 (0.012)	−0.029 (0.011)	−0.023 (0.006)	−0.015 (0.028)	0.008 (0.026)	−0.038 (0.019)
N		334	167	167	334	167	167
Barriers to entrepreneurship	Main effect	−0.013 (0.013)	−0.015 (0.012)	−0.012 (0.007)	−0.049 (0.037)	−0.048 (0.040)	−0.050 (0.0259)
	Interaction	−0.002 (0.012)	−0.003 (0.011)	−0.002 (0.007)	−0.076 (0.033)	−0.094 (0.035)	−0.059 (0.0231)
N		368	184	184	368	184	184
B. Institutions Together							
Labour standards	Main effect	−0.024 (0.016)	−0.030 (0.015)	−0.018 (0.009)	−0.070 (0.067)	−0.049 (0.062)	−0.091 (0.048)
and replacement rate	Labour standards	−0.007 (0.011)	−0.010 (0.010)	−0.003 (0.006)	−0.035 (0.042)	−0.038 (0.038)	−0.033 (0.030)
	Replacement rate	−0.027 (0.012)	−0.030 (0.011)	−0.023 (0.006)	−0.041 (0.035)	−0.024 (0.032)	−0.057 (0.025)
N		334	167	167	334	167	167
Barriers, labour standards,	Main effect	−0.024 (0.016)	−0.030 (0.015)	−0.018 (0.009)	−0.046 (0.053)	−0.019 (0.051)	−0.073 (0.037)
and replacement rate	Barriers	−0.007 (0.023)	0.002 (0.021)	−0.015 (0.013)	−0.093 (0.059)	−0.088 (0.057)	−0.097 (0.041)
	Labour standards	−0.002 (0.020)	−0.011 (0.018)	0.008 (0.011)	0.056 (0.063)	0.051 (0.060)	0.062 (0.044)
	Replacement rate	−0.028 (0.012)	−0.030 (0.011)	−0.026 (0.007)	−0.035 (0.028)	−0.021 (0.027)	−0.049 (0.020)
N		334	167	167	334	167	167

Notes: The Table reports main effects and interaction terms in equation (9) in the text, with the addition of country trends. The instruments and covariates are the same as in Table 6.

5. Summary and Conclusions

This paper presents new evidence on the question of how immigration affects native employment, focusing on the extent to which displacement effects of immigration are mitigated or amplified by cross-country differences in institutions. The estimates typically show that an increase in the foreign share of 10% would reduce native employment rates by 0.2–0.7 of a percentage point. OLS estimates are at the low end of this scale, while the IV estimates using the Balkan Wars 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 identification problems, and the IV estimates are not very precise. It therefore seems reasonable to interpret the OLS and IV estimates as bracketing the true effect.

We are especially interested in the question of whether measures of labour and product-market flexibility change the impact of immigration on native employment. This question has important policy ramifications since many Western European countries are considering institutional and immigration reforms, and working to integrate previous immigrant cohorts more fully. Though restrictive economic institutions can play a protective role for natives, our theoretical framework suggests that institutions such as firing costs, high replacement rates, rigid wages and business entry costs, may ultimately 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 and theoretically more substantial effect that increases the negative impact of immigrants on native employment.

Although not entirely clear cut, the empirical results offer some support for the view that reduced flexibility may make 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 labour 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 for interactions are difficult to identify, the view that restrictive institutions have insulated native workers from competition with immigrants does not get empirical support.

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Data Appendix

Data for Figures 1 and 2

The data plotted in Figure 1 are from OECD (1999*a*) 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 Labour 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% in the Netherlands to 98% in Germany, with the median response rate at 87%. Labour 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) publication *Employment Outlook*.

Time-consistent Definition of Immigrants' EU Status

The analysis here distinguishes non-natives and non-nationals (foreigners) 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. Moreover, we cannot distinguish East from West Germany in the post-unification period. Our analysis uses a time-consistent definition of EU membership. where the EU is defined as the original EU-12 plus Austria and Norway. All immigrants from Germany are also defined as being from the EU (the results are not sensitive to the classification of German immigrants as being from a non-EU country).

Cell Statistics and Confidentiality Edit

At our request, Eurostat provided 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.

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