Revisiting the Effect of Immigration on Native Employment in the EU

Balint Menyhert
Department of Economics, Central European University
Address: 9 Nador utca, 1051 Budapest, Hungary
Phone: +36 1327 3020, fax: +36 1327 3232, e-mail: Menyhert_Balint@ceu-budapest.edu

Abstract

This paper uses and extends the empirical strategy developed by Angrist and Kugler (2003) in order to test the effect of external immigration on native employment rates in the European Union (EU) in light of the booming migration flows of the 2000s. My findings discredit the authors' main assertions that immigration causes considerable job displacement among natives, the extent of which is significantly related to protective labour market institutions in the host country. The divergence of results stems partly from the continuous transformation of the immigration process and its employment consequences for natives, partly from the superior identification strategies employed to account for heterogeneity across countries and effective labour market institutions. Immigration is found to be driven principally by labour demand factors and appears at least as much an economic blessing to entertain as something to be afraid of and protected from.

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1. Introduction

One of the chief immigration concerns for politicians and policy makers alike is the possibility of new entrants depressing labour market opportunities for natives. As expected, the issue has received widespread scholarly attention: the comprehensive meta-analysis carried out by Longhi et al. (2006) documents 165 different estimates of "job robbing" over the last 25 years. The sampled studies are quite varied in terms of both scope and technique, but generally estimate the elasticity of local native employment to immigration to be negative and in the close vicinity of zero. Since most studies with a European focus consider either a single host country or emigrant group, it is very difficult to filter out the systemic effect from the vast array of idiosyncratic and country-specific factors. A notable attempt to overcome this problem is made by Angrist and Kugler (2003) who offer a fresh longitudinal perspective on a Western European panel covering much of the 1980s and 1990s. Not only do these authors tackle potential endogeneity in an ingenious way – by using the Balkan wars as a Mariel boatlift style natural experiment along the lines of Card (1990) –, they also integrate the study of labour market characteristics into the analysis. In particular, Angrist and Kugler find that immigration causes considerable job displacement among natives, the extent of which is significantly related to protective labour market institutions in the host country.

Due to the singularity of this approach in the European context, assessing the external validity of their conclusions is of particular interest. This is all the more important as international migration has exhibited previously unseen patterns and calibre since the turn of the millennium and has induced a great deal of heterogeneity across EU member states. Moreover, the recent financial and economic crisis has also revealed a great deal of information as regards the relationship between migration flows and native employment. Using this as motivation, the present study revisits the issue by updating and extending the empirical strategy developed by Angrist and Kugler (2003). Given the same data source and a similar empirical approach, the results of the two papers are directly comparable.

My results discredit both main assertions of the authors by suggesting that there is no evidence of either negative immigration effect or significant institutional effect on native employment from the early 1990s onwards. The divergence of conclusions is due essentially to three factors. First, the original dataset by Angrist and Kugler (2003) seems to cover a period of transition: partitioning the data into two sub-samples shows that while the posited relationship between immigration and native employment holds for much of the 1980s, it largely disappears during the 1990s. Second, increased immigration flows following the turn of the millennium have affected respective member states in a very different fashion, drawing attention to the problematic nature and required increased dimensionality of the identification problem. More to the point, common estimates can be severely biased due to high levels of heterogeneity in the data panel, an issue not settled appropriately in the original study. Third, exploiting time variation in institutional labour market variables makes it possible to identify the true empirical relationship between immigration and labour market characteristics and discredit associations built on spurious correlations.

The paper is structured as to provide, at every stage, a clear means of comparison with Angrist and Kugler (2003). Section 2 provides some background analysis and presents the statistical properties of the data. Section 3 presents estimates of immigration effects and discusses in detail the identification strategy employed to address potential endogeneity issues. Section 4 considers interaction effects between immigration and protective labour market institutions, while Section 5 concludes.

2. Background and Data

The empirical relationship of interest in this study is the effect of external immigration on native employment rates in Europe. As in the original study, estimations are carried out on a panel of 15 European Union countries and three other Western European economies belonging to the European Economic Area. The main data source is the Eurostat Labour Force Survey, which contains yearly time series of immigration figures and labour market variables for European countries by age, sex, education level, nationality and nativity up to 2010.

Table 1. Descriptive Statistics of Labour Force, Immigrant Shares and Employment Rates for Men (Aged 15-64) in EEA Countries

Country	LFS coverage	Labour force (in millions)			Proportion of non-EU15 non-nationals (%)			Proportion of EU15 non-nationals (%)			Employment-to-population ratio among nationals (%)		
		1989	1999	2009	1989	1999	2009	1989	1999	2009	1989	1999	2009
Austria	1995-2010	2.074	2.160	2.292	-	8.53	8.82	-	1.49	2.60	-	76.34	77.88
Belgium	1983-2010	2.382	2.465	2.579	2.28	3.16	3.69	4.84	7.07	6.14	64.90	68.39	67.71
Denmark	1983-2010	1.527	1.538	1.532	1.16	1.67	3.12	0.79	0.71	1.71	81.30	81.87	78.66
Finland	1995-2010	1.366	1.345	1.359	-	1.00	1.78	-	0.32	0.57	-	70.47	69.77
France	1983-2010	14.137	14.304	15.163	3.72	4.64	3.85	2.94	2.50	2.11	72.20	68.39	69.07
Germany	1983-2010	21.842	22.697	22.747	5.01	7.01	7.25	3.22	3.02	3.08	71.60	73.34	76.34
Greece	1983-2010	2.601	2.887	3.037	0.51	3.75	10.12	0.11	0.16	0.21	64.20	71.18	72.61
Iceland	1995-2010	0.720	0.820	0.940	-	-	3.39	-	-	1.30	-	89.58	80.25
Ireland	1983-2010	0.865	0.985	1.169	0.47	0.39	11.11	1.77	2.82	4.46	56.60	74.08	65.70
Italy	1992-2010	14.898	14.072	14.741	-	0.86	8.25	0.06	0.13	0.21	-	60.30	67.71
Luxemburg	1983-2010	0.104	0.111	0.131	2.26	4.15	5.56	30.00	35.70	43.60	64.00	73.34	70.47
Netherlands	1983-2010	4.141	4.495	4.800	2.19	1.86	2.14	1.67	1.80	1.51	66.30	81.87	82.70
Norway	1995-2010	1.155	1.242	1.337	-	1.35	3.02	-	1.67	2.76		75.58	78.66
Portugal	1986-2010	2.627	2.687	2.781	0.59	0.92	4.26	0.22	0.34	0.38	72.90	81.87	71.18
Spain	1986-2010	10.303	10.718	12.820	0.15	1.37	14.10	0.33	0.62	1.62	55.60	69.07	67.71
Sweden	1995-2010	2.393	2.304	2.521	-	2.32	3.03	-	1.59	2.31	-	73.34	74.83
Switzerland	1996-2010	2.077	2.153	2.289	-	5.16	8.89	-	17.38	17.03	-	88.69	84.37
United Kingdom	1983-2010	16.287	15.758	16.691	2.94	2.29	6.66	1.77	1.54	1.80	76.50	77.11	74.83

Source: Eurostat Labour Force Survey

Descriptive statistics presented in Table 1 list the sample countries and show the evolution of external ('non-EU15') and internal ('EU15') immigrant shares as well as native employment-to-population ratios for men during the last two decades. Even though the figures contain a great deal of variation in both dimensions, three patterns clearly stand out. First, immigration from within the EU has remained essentially unaffected over the years.

Second, the sample covers two markedly different sub-periods with respect to external immigration: while fairly stable shares characterised the 1990s, the 2000s saw exuberant increases in this regard. Third, the bulk of the abundant inflow of non-EU15 immigrants has concerned only a handful of countries (most notably Greece, Ireland, Italy, Portugal and Spain), the ones exhibiting the largest variations in employment rates as well. Not contained in these figures but nevertheless an empirical fact, the onset of the current crisis has brought on stark reductions in external immigration shares, in the most heavily affected countries in particular.

Following the Eurostat classification also used by Angrist and Kugler, I measure immigrant shares in each country as the proportion of the active labour force that is not a national of any of the then 15 listed member states of the European Union (prior to the 2004 and 2007 rounds of enlargements). Employment rates for natives, on the other hand, are relative to the entire population within a given age group.

3. Estimation of Immigration Effect

3.1. OLS Estimates

The baseline estimates of immigration effect are obtained by panel regressions allowing for a single immigration effect for each country. The estimated model is given by

$$ln(y_{ijt}) = \mu_i + \alpha_j + \delta_t + \beta_i \cdot ln(s_{ijt}) + \varepsilon_{ijt} \quad (1)$$

where subscripts i, j and t denote demographic groups, countries and years, respectively. This specification thus includes fixed group, country and year effects and associates the main parameter with each demographic group. The regressor $\ln(s_{ijt})$ is the log share of active immigrants, while the dependent variable $\ln(y_{itj})$ is the log share of employment-to-population ratio for natives. Following Angrist and Kugler (2003), immigrant share is defined as the proportion of labour force not national of any state in the EU15 area.

The proposed model relies on time-invariant country effects and the absence of omitted factors correlated with immigration that exert a direct influence on natives' employment rate. The most apparent omissions are time-varying productivity or labour demand shocks correlated with both immigrant shares and native employment. It is therefore opportune to include controls for the log share of active foreigners with EU15 nationality: not only should this provide a partial (though potentially endogenous) control for local demand factors that may influence immigration, but it also addresses the point, raised in earlier immigration studies such as Card (2001), that internal migration flows can partially offset exogenous changes in the labour force. Also, estimates in Equation (1) are biased if immigration is correlated with country-specific employment trends, a problem made more likely by the length of the sample period. To this effect, extended models are also considered that replace the fixed country effect α_j with a country-specific (linear) trend $\alpha_{0j} + \alpha_{1j} \cdot t$.

Angrist and Kugler (2003) obtain insignificant OLS estimates for native employment effect for both sexes, except for the small and significant effect of -0.021 when the sample is limited to young men. Including the EU15 immigration share as a control does not qualitatively change the results, although it leads to an even larger negative estimate of -0.037 for young men and a significant effect of -0.021 for men overall. These results are obtained on the full sample and change markedly once the sub-periods of 1983-1990 and 1991-1999 are analysed separately. In addition, this partition allows for a direct comparison with my own estimates.

Table 2. OLS Estimates of the Main Effect of Immigration on Native Employment Rates

	ANGRIST AND KUGLER ESTIMATES			OWN ESTIMATES			
	Overall	By age group		Overall	By age group		
		Under 40	Over 40		Under 25	25-54	Over 55
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
A. 1983 - 1990							
Without trend							
Men	-0.047* (0.020)	-0.065* (0.020)	-0.029 (0.016)				
Women	-0.059 (0.131)	-0.097* (0.032)	-0.022 (0.039)				
With trend							
Men	-0.032 (0.021)	-0.030* (0.015)	-0.034* (0.012)				
Women	-0.041 (0.177)	-0.034* (0.016)	-0.046 (0.031)				
B. 1991 - 1999							
Without trend							
Men	-0.004 (0.013)	-0.016 (0.015)	0.007 (0.010)	0.032 (0.032)	0.123* (0.049)	0.040 (0.022)	-0.009 (0.021)
Women	0.017 (0.029)	0.010 (0.018)	0.024 (0.021)	0.064* (0.029)	-0.047 (0.090)	0.052 (0.029)	-0.115 (0.110)
With trend							
Men	-0.006 (0.022)	-0.008 (0.016)	-0.004 (0.010)	-0.010 (0.013)	0.016 (0.033)	0.000 (0.011)	-0.001 (0.020)
Women	-0.009 (0.051)	-0.004 (0.013)	-0.013 (0.014)	0.009 (0.024)	-0.006 (0.023)	0.017 (0.025)	-0.012 (0.037)
C. 2000 - 2010							
Without trend							
Men				-0.016 (0.010)	-0.038 (0.021)	-0.021* (0.006)	-0.054* (0.019)
Women				0.022 (0.024)	0.019 (0.028)	0.021 (0.025)	-0.045 (0.061)
With trend							
Men				0.072* (0.020)	0.129* (0.047)	0.057* (0.017)	0.031 (0.024)
Women				0.048* (0.015)	0.087* (0.030)	0.041* (0.015)	-0.013 (0.023)

Note: The table reports OLS estimates of β_i in Equation (1). Robust standard errors are reported in parenthesis. Statistically significant estimates at 5% per cent probability level are marked with an asterisk.
Source: Author's calculations

Panels A, B and C of Table 2 show results for each of the last three decades, separately for men and women, both with and without country specific trends. Columns 1-3 report estimates obtained on the dataset by Angrist and Kugler (2003), while columns 4-7 contain my own results. The figures indicate a very interesting pattern: while immigration was associated with significant native job losses in the 1980s, this negative effect generally disappeared during the 1990s. Moreover, once country-specific employment trends are accounted for, strong positive relationship prevailed between immigration and native employment for almost all demographic groups during the 2000s. The magnitude of the estimated elasticities might be illustrated with a numerical example: in a country where 5%

of the labour force is of non-EU15 nationality, the arrival of 100 active immigrants would destroy about 70 to 100 native jobs during the 1980s – an almost one-to-one job displacement. After 2000, however, the same exogenous change would be met with the creation of 80-150 additional native jobs.

Reverse causality may have something to do with the major turnaround in the observed effect of immigration over the years. Standard economic theory offers two straightforward explanations: either the European labour market has become more integrated over the years so that increases in labour demand has affected native and immigrant workers in a more and more similar fashion or, alternatively, the rise in immigration has induced adjustments in factor prices that has increased the labour supply of natives. In theory, the use of immigration shares from within the EU15 area to control for labour demand factors, as proposed by Angrist and Kugler (2003), should help in resolving the question. The fact that the additional control turns out to be insignificant most of the time and exerts only negligible influence on the presented estimates seems to imply that it is either an unsatisfactory measure of demand or that the business cycle concerns mostly external immigrants. In the next section, alternative approaches are considered.

3.2. Instrumental Variable Estimates

The ordinary least squares (OLS) estimates in Table 1 may be biased upwards if individuals choose to migrate to countries where and when their employment prospects are best. This section discusses estimates of immigration effects using different instrumental variable (IV) strategies. The choice of instruments is motivated by Angrist and Kugler (2003) in an attempt to offer direct means of comparison with their findings.

It is worthwhile to start with the Angrist and Kugler dataset and implement their IV strategy on the separate subsamples of the 1980s and 1990s. As a reminder, in their original study Angrist and Kugler, motivated by a sharp run-up in the number of Yugoslavs among European immigrants in the early and late 1990s due to wars in Bosnia and Kosovo, chose the distance from the former Yugoslav republics as a predictor of immigrant shares. Their reported IV estimates are generally significant and larger in magnitude than the corresponding OLS estimates, although they imply an implausibly high (more than one-to-one) job displacement, especially among women. While using this IV strategy beyond the 1990s is not adequate due to both large drops in Yugoslav immigration and the enormous subsequent growth of international migration, it is nevertheless instructive to see how the instruments perform in the 1991-1999 period relative to the full sample. Columns 1-3 of Table 3 show Angrist and Kugler's IV estimates on the full sample as reported in their paper, while columns 4-6 contain my own estimates on the same data for the 1990s.o Results show that the identification strategy based on Yugoslav immigration ceases to yield negative and significant estimates for immigration effect once applied solely to the war-torn 1991-1999 period. This means that there are no war-related exogenous breaks identified in the time-series behavior of employment rates for countries relatively close to Yugoslavia after the outbreak of the Bosnian war. The more pronounced IV estimates reported by Angrist and Kugler might therefore be caused by confusing the simultaneous employment effects of economic depression (due to the collapse of the Soviet Union and adherent socialist economies) and increased immigration (due to the Yugoslav conflict) in Central Europe during the early 1990s.

Table 3. IV Estimates for the Main Effect Using the Identification Strategy of Angrist and Kugler (2003) on the Original Dataset

	FULL SAMPLE (1983 - 1999)			SECOND SUBSAMPLE (1991 - 1999)		
	Overall	By age group		Overall	By age group	
		Under 40	Over 40		Under 40	Over 40
	(1)	(2)	(3)	(4)	(5)	(6)
Men	-0.050* (0.023)	-0.082* (0.030)	-0.018 (0.016)	-0.003 (0.020)	-0.006 (0.025)	-0.001 (0.016)
Women	-0.245* (0.093)	-0.189* (0.070)	-0.301* (0.102)	0.007 (0.035)	0.042* (0.018)	-0.027 (0.022)

Note: The table reports 2SLS estimates of β_i in Equation (1) obtained on the original dataset when the regressor is instrumented for by the method described in Angrist and Kugler (2003). Robust standard errors are reported in parenthesis. Statistically significant estimates at 5% per cent probability level are marked with an asterisk.
Source: Author's calculations

Reliable assessment of immigration effects after the millennium, however, requires an IV strategy that exploits more recent migration trends. The fact that the overwhelming majority of immigrants to the EU15 countries is young and comes from Eastern Europe, the wider Balkans and the Maghreb countries suggests that youth unemployment in these regions should be a good predictor of the foreign immigrant shares. Therefore, in close correspondence to Angrist and Kugler (2003), I equally make exogenous changes in the excess supply of labour on the periphery of Western Europe the backbone of my identification strategy and apply the distance between origin and destination countries as the main driver of migration flows. To this effect, I create time-series of weighted averages of yearly youth unemployment for Eastern Europe, South-Eastern Europe and North Africa using the Key Indicators of the Labour Market database of the International Labour Organisation.[1] Also, I measure the distance between each EU15 capital and Warsaw (Poland), Istanbul (Turkey) and Tunis (Tunisia) as the chosen geographical and population centres of the respective emigration regions. By taking the cross-products and reweighting outer unemployment rates by relative distance, I claim to capture in an econometrically meaningful way the interaction of youth unemployment abroad and geographical proximity as the important determinants of immigration flows.

The first-stage equation for the IV estimates for a given demographic group is given by

where $$ln(s_{jt}) = \phi_j + \psi_t + \gamma_{EE} \cdot q_{jt}^{EE} + \gamma_{SE} \cdot q_{jt}^{SE} + \gamma_{NA} \cdot q_{jt}^{NA} + \xi_{jt} \qquad (2)$$

q_{jt}^{EE} = distance from Warsaw × average youth unemployment rate in Eastern Europe

q_{jt}^{SE} = distance from Istanbul × average youth unemployment rate in Southeast Europe

q_{jt}^{NA} = distance from Tunis × average youth unemployment rate in North Africa

are the excluded instruments, while ϕ_j and ψ_t are country and year effects. Distance from potential host countries is measured by the inverse of normalised geographic distance,

[1] The respective regions are made to consist of the following countries:
- Eastern Europe: The Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia;
- South-Eastern Europe: Albania, Bosnia Herzegovina, Bulgaria, Croatia, Macedonia, Montenegro, Romania, Serbia, Turkey;
- North Africa: Algeria, Egypt, Libya, Morocco, Tunisia.

whereby a proportionate adjustment to youth unemployment rates is applied. Alternative specifications with no year effects but country-specific coefficients $\gamma_{j,EE}$, $\gamma_{j,SE}$ and $\gamma_{j,NA}$ are also considered, in an attempt to account for institutional patterns or diaspora effects. In general, first-stage coefficients for Eastern and South-Eastern Europe turn out to be strongly significant and indicative of a positive relationship between regressors and immigration shares on the whole, although the coefficients for North Africa remain insignificant throughout. To ascertain the validity of the proposed identification strategy, I have also considered models in which non-EU foreign share is replaced within EU15 foreign share as the dependent variable. The resulting estimates show no relationship between (distance-weighted) youth unemployment on the European periphery and internal EU15 migration flows, which confirms that instruments are indeed exogenous and assumed to influence native employment rates in EU15 countries only through their effect on migration.

The two-stage least squares (2SLS) estimates for the main immigration effect using q_{jt}^{EE}, q_{jt}^{SE} and q_{jt}^{NE} as instruments are reported in Table 3 for models both with and without country-specific employment trends on the most recent sub-sample covering the 2000s. Columns 1-4 show estimates obtained by common parameters γ_{EE}, γ_{SE} and γ_{NA} in the first stage, while columns 5-8 contain estimates derived from allowing first-stage coefficients to vary across countries. Results indicate that the respective first stage specifications matter very little when it comes to driving main estimates of the employment effect of immigration. This is not the case with the second stage equation: while the baseline specification without country-specific employment trends yield negative and generally significant estimates for men (as shown in Panel A), the inclusion of trends essentially reverses the sign of these estimates without compromising their statistical significance (as shown in Panel B). As estimates for women are in general not significantly different from zero, 2SLS results are qualitatively identical to the respective OLS estimates presented in Panel C of Table 4 for all demographic groups.

Table 4. IV Estimates for the Main Effect for the 2000s

	BASELINE FIRST-STAGE MODEL				COUNTRY-SPECIFIC FIRST-STAGE MODEL			
	Overall	By age group			Overall	By age group		
		Under 25	25-54	Over 55		Under 25	25-54	Over 55
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Without country-specific trend								
Men	-0.100* (0.014)	-0.050 (0.042)	-0.037* (0.017)	-0.031 (0.038)	-0.031* (0.008)	-0.049 (0.028)	-0.026* (0.008)	-0.050* (0.016)
Women	-0.049 (0.064)	-0.011 (0.036)	0.044 (0.046)	0.193* (0.095)	0.021 (0.025)	-0.001 (0.018)	0.033 (0.030)	-0.072 (0.060)
B. With country-specific trend								
Men	-0.019 (0.018)	0.074* (0.031)	0.018* (0.006)	0.036* (0.013)	0.031 (0.019)	0.046 (0.051)	0.041* (0.010)	0.057* (0.025)
Women	-0.014 (0.017)	0.035 (0.039)	0.001 (0.008)	-0.022 (0.043)	0.031* (0.012)	0.041 (0.046)	0.027 (0.008)	0.012 (0.022)

Notes: The table reports 2SLS estimates of β_i in Equation (1) obtained by different first-stage specifications. Robust standard errors are reported in parenthesis. Statistically significant estimates at 5% per cent probability level are marked with an asterisk.
Source: Author's calculations

IV estimates thus imply considerable immigration-driven job displacement or job creation depending on the main specification. To determine the true effect of immigration on native employment within these extremes, some additional identification is needed. This may come in the form of estimating the employment consequences of immigration on more homogeneous groups of countries, a strategy made all the more warranted by the divergent 2SLS estimates across different specifications: since year dummies capture systematic changes in employment rates, allowing for country-specific trends should have an effect only if employment histories are considerably different between member states. By focusing on countries that exhibit more similar employment and immigration patterns, not only are estimates expected to converge across specifications, the robustness of the main effect can also be assessed.

For this end, I looked at the immigration and employment history of each country in the sample and created two markedly different but fairly homogeneous sets. The first group, consisting of Finland, Greece, Ireland, Portugal and Spain, is characterized by high exposure to immigration from outside the EU15: despite low initial immigration levels at the beginning of the period, these countries have experienced the highest increases and variation in immigration shares over the last decade, coupled with a high of sensitivity of native employment rates. The second group, consisting of Belgium, France, Germany, the Netherlands and Switzerland, is the exact opposite: albeit having high absolute immigration levels, countries belonging here have faced the lowest change and variation in immigrant shares and employment rates over the sample period.

Table 5. IV Estimates for the Main Effect Across Country Groups Affected Differently by Immigration

	HIGH EXPOSURE TO IMMIGRATION				LOW EXPOSURE TO IMMIGRATION			
	Overall	By age group			Overall	By age group		
		Under 25	25-54	Over 55		Under 25	25-54	Over 55
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
A. Without country-specific trend								
Men	0.067 (0.058)	0.009 (0.013)	-0.006 (0.008)	-0.046 (0.028)	-0.023 (0.087)	-0.079 (0.033)	-0.029 (0.011)	-0.111* (0.024)
Women	0.042 (0.065)	0.040 (0.021)	0.108* (0.009)	0.084 (0.055)	-0.042 (0.090)	0.004 (0.007)	-0.007 (0.006)	-0.364* (0.102)
B. With country-specific trend								
Men	-0.021 (0.039)	0.028 (0.020)	0.013 (0.007)	-0.008 (0.018)	-0.013 (0.046)	0.039 (0.040)	0.010 (0.017)	0.006 (0.025)
Women	0.012 (0.013)	0.041 (0.018)	0.009* (0.001)	-0.013 (0.032)	0.029 (0.042)	0.036 (0.039)	-0.012 (0.015)	0.009 (0.008)

Notes: The table reports 2SLS estimates of β_i in Equation (1) obtained for different country clusters with the baseline first-stage model featuring common interaction parameters. Robust standard errors are reported in parenthesis. Statistically significant estimates at 5% per cent probability level are marked with an asterisk.
Source: Author's calculations

2SLS results for high and low exposure countries are shown separately in columns 1-4 and columns 5-8 of Table 5. In general, estimates for respective demographic groups obtained with and without country-specific trends in the main equation become statistically indistinguishable from one another. Moreover, estimated coefficients turn out to be insignificant on the whole, regardless of the country group under consideration. This implies that no evidence of native job displacement due to immigration is found in the data from the 1990s onwards. Apparently, EU15 member states have been able to accommodate immigration flows of great magnitude also, most probably through economic expansion. The hypothesis that labour demand factors are the principal drivers of immigration is further corroborated by the steep decline in the stock of immigrants documented in member states following the onset of the recent financial and economic crisis.

4. Interaction of Immigration with Labour Market Institutions

The second main assertion of Angrist and Kugler (2003) is that labour market institutions can change the employment consequences of immigration for natives. In particular, restrictive or rigid labour standards that apply differently to natives and foreigners seem to aggravate immigration-fuelled job displacement. Applying and extending the canonical two-factor labour demand model of Hamermesh (1986), Angrist and Kugler (2003) conjecture that higher firing costs and replacement rates increase job destruction among natives by lowering the relative cost of employing foreigners and improving natives' non-work options, respectively. Wage flexibility and low barriers to business entry, on the other hand, are assumed to help job creation and act to offset the negative employment effects of immigration on natives. By accounting for institutional characteristics in a consistent way, the authors claim to have given a more appropriate and subtle description of the labour market processes set forth by immigration.

As for the empirical strategy, Angrist and Kugler (2003) use three different measures of institutional features. The first is a summary measure of labour standards based on the extent of employment protection as well as restrictions on work hours and employment contracts taken from OECD data compiled by Nickell (1997). The second measure is the average replacement rate from the same source, while the third is an index of barriers to entrepreneurship assembled by Nicoletti et al. (2000). The equation used for estimating interactions between immigration and labour market institutions is

$$ln(y_{ijt}) = \mu_i + \delta_t + \alpha_j + (\beta_{0i} + \beta_{1i} \cdot x_j) ln(s_{ijt}) + \varepsilon_{ijt} \tag{3}$$

where x_j is a normalised country-specific institutional variable. The parameter β_{0i} therefore captures the main effect of immigration on demographic group i in countries with average institutional features, while the interaction term β_{1i} describes how this effect changes with each standard deviation change in x_j. On the full sample for men, Angrist and Kugler (2003) find that immigration effects are more negative in countries with less flexible labour markets, higher replacement rates and entry costs. OLS estimates of the interaction effect are significant around -0.020 regardless of the institutional variable, while the main effect increases somewhat in both significance and magnitude. Instrument-based 2SLS estimates in general further amplify both the main and the interaction effect, even though the use of the replacement rate measure leads to insignificance of both.

Similarly to the initial test applied with respect to the baseline results, it is instructive to start with partitioning the original data into two distinct sub-periods and estimate employment effects separately using Equation (3). Table 6 shows OLS estimates for both the 1980s and 1990s and 2SLS for the 1990s, thereby offering a direct means of comparison with results presented in Table 6 of Angrist and Kugler (2003). Figures in columns 1-3 indicate that while both the main and interaction estimates are typically negative, borderline significant and remain on the whole similar to the original ones for the 1980s, they as a rule lose their statistical significance in the 1990s sample.

Table 6. OLS and 2SLS Estimates of Main and Labour Market Interaction Effects of Immigration for Men Using the Identification Strategy of Angrist and Kugler (2003) on the Original Dataset

	OLS ESTIMATES			2SLS ESTIMATES		
	(1)	(2)	(3)	(4)	(5)	(6)
	Overall	By age group		Overall	By age group	
		Under 40	Over 40		Under 40	Over 40
A. 1983 - 1990						
Labour standards						
Main effect	-0.098 (0.057)	-0.113* (0.050)	-0.082* (0.036)			
Interaction effect	-0.019 (0.028)	-0.023 (0.026)	-0.016 (0.016)			
Replacement rate						
Main effect	-0.079* (0.022)	-0.092* (0.023)	-0.068* (0.022)			
Interaction effect	-0.034* (0.017)	-0.043* (0.018)	-0.024 (0.018)			
Barriers to entrepreneurship						
Main effect	-0.039 (0.021)	-0.067* (0.014)	-0.034* (0.016)			
Interaction effect	0.003* (0.001)	0.028 (0.019)	0.038* (0.016)			
B. 1991 - 1999						
Labour standards						
Main effect	-0.004 (0.020)	-0.016 (0.024)	0.009 (0.014)	0.113* (0.043)	0.136* (0.037)	0.090* (0.039)
Interaction effect	-0.024 (0.016)	-0.028 (0.022)	-0.020 (0.011)	-0.064* (0.020)	-0.091* (0.023)	-0.038* (0.019)
Replacement rate						
Main effect	-0.013 (0.022)	-0.028 (0.022)	0.001 (0.014)	0.121* (0.047)	0.132* (0.047)	0.111* (0.038)
Interaction effect	-0.010 (0.013)	-0.014 (0.015)	-0.006 (0.009)	0.027 (0.049)	0.027 (0.023)	0.028 (0.015)
Barriers to entrepreneurship						
Main effect	-0.011 (0.017)	-0.036* (0.017)	0.004 (0.010)	-0.012 (0.023)	-0.022 (0.028)	-0.005 (0.016)
Interaction effect	0.004* (0.001)	0.005* (0.002)	-0.038* (0.013)	0.004* (0.001)	0.005* (0.002)	-0.040* (0.014)

Note: The table reports OLS and 2SLS estimates of β_{0i} and β_{1i} in Equation (3) augmented by country-specific trends for the 2000-2010 period for men. Robust standard errors are reported in parenthesis. Statistically significant estimates at 5% per cent probability level are marked with an asterisk.
Source: Author's calculations

To assess the performance of institutional variables in explaining native employment patterns more recently, regressions are re-run for the 2000s both with OLS and 2SLS based on the identification strategy put forward in the previous section. Furthermore, I extend the analysis by accounting for labour market institutions in a more comprehensive and nuanced manner than in the original study. Besides employing a single cross-section of the respective institutional variable x_j as a proxy for labour market institutions, I also exploit observed time variation in the institutional regressors. Specifically, I use the time series of OECD synthetic indicator of employment protection to capture labour standards and the time series of OECD summary measure of benefit entitlements to account for changes in the replacement rate. As for measuring barriers to entrepreneurship, in absence of sufficient OECD data as used by Angrist and Kugler (2003), I use the comprehensive Index of Economic Freedom assembled by the Heritage Foundation for this purpose.[2] This way, it becomes possible to capture changes in labour market institutions over time and ascertain whether employment outcomes are effectively driven by them or not. As a benchmark, I also consider estimates based on the country-specific means of the institutional variable series in order to provide comparability with the original estimates.

Table 7. OLS and 2SLS Estimates of Main and Labour Market Interaction Effects with Different Institutional Variables for the 2000s

	OLS ESTIMATES				2SLS ESTIMATES			
	Overall	By age group			Overall	By age group		
		Under 25	25-54	Over 55		Under 25	25-54	Over 55
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Static labour market institutions								
Labour standards								
Main effect	0.074* (0.018)	0.015 (0.098)	0.053 (0.026)	0.008 (0.046)	-0.026 (0.025)	0.081 (0.060)	0.021 (0.021)	0.065 (0.034)
Interaction effect	0.011 (0.019)	0.202 (0.133)	0.010 (0.023)	0.042 (0.059)	-0.008 (0.031)	-0.009 (0.049)	-0.002 (0.016)	-0.032 (0.033)
Replacement rate								
Main effect	0.075* (0.019)	0.078 (0.039)	0.046 (0.013)	0.010 (0.018)	-0.022 (0.016)	0.035 (0.055)	0.020 (0.015)	0.053 (0.024)
Interaction effect	-0.016 (0.021)	-0.178* (0.036)	-0.045* (0.011)	-0.079* (0.017)	0.031 (0.017)	-0.064* (0.022)	0.002 (0.011)	0.029 (0.023)
Barriers to entrepreneurship								
Main effect	0.078* (0.025)	0.109 (0.057)	0.056* (0.017)	0.027 (0.022)	-0.027 (0.013)	0.072 (0.047)	0.020 (0.015)	0.044 (0.022)
Interaction effect	-0.009 (0.026)	-0.083 (0.131)	-0.015 (0.022)	-0.034 (0.023)	0.034 (0.034)	-0.005 (0.061)	0.004 (0.021)	0.029 (0.037)
B. Dynamic labour market institutions								
Labour standards								
Main effect	0.028* (0.012)	-0.007 (0.015)	0.015 (0.009)	-0.034 (0.029)	-0.072* (0.015)	0.045 (0.025)	0.014 (0.016)	0.046 (0.023)
Interaction effect	0.003 (0.005)	0.009 (0.016)	0.002 (0.002)	-0.006 (0.014)	0.004 (0.002)	-0.015 (0.008)	-0.002 (0.004)	-0.010 (0.013)
Replacement rate								
Main effect	0.025 (0.013)	-0.000 (0.018)	0.018 (0.010)	-0.027 (0.025)	-0.065* (0.018)	0.039 (0.021)	0.011 (0.013)	0.037 (0.021)
Interaction effect	-0.004* (0.001)	-0.007 (0.007)	-0.002 (0.003)	-0.009 (0.005)	-0.002 (0.003)	0.001 (0.007)	-0.001 (0.002)	0.001 (0.006)
Barriers to entrepreneurship								
Main effect	0.073* (0.019)	0.130* (0.045)	0.058* (0.016)	0.034 (0.023)	-0.019 (0.017)	0.077 (0.037)	0.019 (0.007)	0.037 (0.014)
Interaction effect	-0.001 (0.002)	-0.002 (0.019)	-0.002 (0.002)	-0.004 (0.002)	-0.002 (0.003)	-0.015 (0.009)	-0.002 (0.001)	-0.004 (0.004)

Note: The table reports OLS and 2SLS estimates of β_{0i} and β_{1i} in Equation (3) augmented by country-specific trends for the 2000-2010 period for men, both with static and dynamic institutions in the generation of interaction variables. Robust standard errors are reported in parenthesis. Statistically significant estimates at 5% per cent probability level are marked with an asterisk.
Source: Author's calculations

[2] The respective data are available under the following URLs:
http://www.oecd.org/document/34/0,3746,en_2649_33927_40917154_1_1_1_1,00.html#epl
http://www.heritage.org/index/explore

Table 7 presents respective OLS and 2SLS estimates of the main and interaction effects for men derived from the trend-augmented model on the 2000s. Panel A contains estimates associated with a single cross-section of institutional variables, while Panel B shows results derived from the richer representation of labour market institutions. With static institutional variables, estimates for the interaction effect are qualitatively similar to the ones presented by Angrist and Kugler (2003) in the sense of being mostly negative and comparable in magnitude to the main effect. They generally turn out to be not significantly different from zero, although the (implausibly high) negative interaction effect obtained with the replacement rate measure seems to offer some scarce evidence in favour of the authors' proposition. Estimates based on dynamic institutional variables (presented in Panel B of Table 7), however, clearly go against this hypothesis as respective interaction estimates turn out to be insignificant throughout and typically an order of magnitude smaller than with static institutions. As for the main estimates, they are comparable to baseline OLS and 2SLS estimates (presented in Panel C of Table 2 and Panel B of Table 4, respectively) for all labour market measures, even though OLS estimates become more ambiguous and 2SLS estimates for men overall appear more negative and tend to gain in significance. This implies that no legitimate evidence is found over the last two decades to maintain that labour market institutions in EU15 countries exert considerable influence on the way immigration affects native employment.

Conclusions

This paper uses and extends the empirical strategy developed by Angrist and Kugler (2003) in order to test the effect of external immigration on native employment rates in the European Union. These authors find that immigration causes considerable job displacement among natives, the extent of which is significantly related to protective labour market institutions in the host country. Using the same data source and a comparable empirical approach, this study aims to revisit the issue and update estimates in light of the booming migration flows of the 2000s and the recent economic crisis.

My results discredit both main assertions of the original paper as I find no evidence of either the negative native employment consequences of immigration or the strong influence of labour market institutions on the degree of job displacement since the early 1990s. This divergence from the results and conclusions of Angrist and Kugler (2003) stems mainly from three sources. First, the original dataset lumps together two distinct periods: partitioning the data into two sub-samples shows that while the posited relationship between immigration and native employment seems indeed to hold for much of the 1980s, it largely disappears during the 1990s. Second, the turn of the millennium has brought about a previously unseen amplification of immigration to certain countries, which resulted in a great heterogeneity across countries in the sample. The proper identification of the true immigration effect thus requires accounting for the increased dimensionality of the phenomenon at hand. Third, by exploiting time variation in institutional labour market variables, it becomes possible to decipher true connections between immigration and labour market characteristics from other possible associations built on correlations with one or the another.

Findings presented in this paper seem to imply that immigration is driven principally by labour demand factors, as suggested by the empirical refutation of the job displacement hypothesis and evidenced by the stark decrease in immigration shares after the onset of the recent crisis. Far from presenting unassailable evidence on the subject, my results nevertheless hint that immigration is at least as much an economic blessing to entertain as it is something to be afraid of and protected from.

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