Migration and Wage-Setting: Reassessing the Labor Market Effects of Migration

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Abstract: This paper employs a wage-setting approach to analyze the labor market effects of immigration into Germany from 1980 to 2004. This enables us to consider labor market rigidities, which are prevalent in Europe. We find that the elasticity of the wage-setting curve is particularly high for young workers. Moreover, natives and foreigners are imperfect substitutes. The wage and employment effects of immigration depend on the skill structure of the immigrant workforce. Since the foreign labor supply shift mainly affected the high-skilled labor market segment, the four percent increase of the workforce through immigration did neither increase aggregate nor foreign unemployment.

Keywords: immigration, wages, unemployment, Germany, panel data

JEL code: F22, J31, J61.

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I. Introduction

High and increasing immigration rates in the US and Europe have fanned fears that migrants reduce wages and limit the employment opportunities of the native labor force. In continental European countries – where labor market rigidities are prevalent and unemployment is already high and persistent – concerns that immigration will exacerbate unemployment are particularly widespread. Moreover, disproportionately high unemployment rates among the foreign workforce have contributed to popular perceptions that further immigration will create an unsustainable fiscal burden on European welfare states. If immigration does indeed have adverse effects on the labor market, continental European governments may come under further pressure to tighten their immigration policies.

In this paper we present a new approach to measuring the wage and employment effects of migration under the conditions of imperfect labor markets. With this approach, we are able to address the effects of migration on unemployment in greater depth than earlier studies have done. Our approach is based on a wage-setting framework in which we assume that wages tend to decline with increasing unemployment under given price expectations, albeit imperfectly. Collective bargaining and efficiency wage models suggest that the elasticity between wages and the unemployment rate varies between groups in the labor market (Layard et al. 2005). In our empirical analysis we therefore estimate wage-setting curves that, together with the elasticities of labor demand, determine the wage and employment effects of immigration in different segments of the labor market. This enables us to take into account labor market rigidities, which exercise a particularly strong impact in the European context, affecting different types of labor in different ways.

Our empirical analysis focuses on Germany, the third most popular destination for migrants in the world after the US and Russia (Freeman 2006). Following the 1989 collapse of the Communist regimes of Central and Eastern Europe, the net immigration rate in Western Germany climbed from about zero in the early 1980s to about six per thousand at the beginning of the 1990s, compared to three per thousand in the fifteen member states of the European Union (EU-15) and four per thousand in the US (World Bank 2007). However, since the beginning of this millennium, the net immigration rate has dropped to less than three per thousand in the wake of Germany's economic downturn.

Germany is also an interesting case because it has been suffering from high and persistent unemployment since the first oil price shock in 1973. The unemployment rate has increased further since German unification in 1990. Moreover, due to Germany’s extensive labor market regulations, the labor market is considered to be highly rigid (OECD 1994), and collective bargaining still plays an import role in wage-setting. About 85 percent of employees are directly or indirectly covered by collective agreements negotiated mainly at the industry level (Ellguth/Kohaut 2007). However, efficiency wages play an important role as well. About 46 percent of the firms bound by collective agreements pay wages above the levels stipulated in the agreements (Jung/Schnabel 2009).

While our approach derives the labor market effects of immigration from an equilibrium framework, the overwhelming majority of the empirical literature estimates reduced-form equations that relate wages or employment variables to the immigrant share in specific geographic areas or industries (Card 1990, Hunt 1992, Pischke/Velling 1997, Dustmann et al. 2005). These studies find only very small wage and employment effects, if any at all (Longhi et al. 2006, 2008). The spatial correlation approach has been criticized, however, for yielding spurious results if immigrants are not randomly distributed across local labor markets (or industries) or if other factors equilibrate labor market conditions across geographical areas (Borjas 2003). Some recent studies have therefore used the variance in immigrant shares across education and experience groups at the national level, assuming that the allocation of immigrants across skill and experience groups is exogenous (Borjas 2003, Aydemir/Borjas 2006). Nevertheless, this literature estimates only partial correlations between wage or employment variables on the one hand and the immigrant share on the other, and does not consider the interaction between wages and employment and the cross-effects of labor supply shifts in different segments of the labor market.

Another strand of the empirical literature derives the wage effects of migration from an aggregate production function approach (Grossmann 1992, Borjas 2003, Ottaviano/Peri 2006, 2008, Manacorda et al. 2006, Dustmann et al. 2008). This general equilibrium approach takes cross-effects of labor supply shifts on wages in different segments of the labor market into account, but it relies on the assumptions of perfect competition and clearing labor markets and thus cannot address the effects of migration on wages and employment if labor markets are imperfect.

Closer to our approach are some recent studies that analyze the effects of immigration on the Phillips curve. Binyamini/Razin (2008) and Engler (2007) show that migration can alter the elasticity of labor demand and supply, inducing a flatter Phillips curve. Similarly, Bentolila et al. (2008) find that immigration has reduced the wage elasticity of labor supply and the wage-markup in Spain, which in turn has weakened the trade-off between inflation and unemployment there. The present paper differs from these contributions in that it focuses on the long-run equilibrium relationship between wages and unemployment and does not address the impact of immigration on the inflation-unemployment trade-off. Moreover, we consider disaggregated labor supply shifts in different education and experience segments of the labor market rather than aggregate labor supply shocks.

Three recent papers that address the wage and employment effects of immigration in Germany are related to our study: Glitz (2006) examines the labor market effects of the large-scale immigration of ethnic Germans (so-called “Spätaussiedler”) based on the spatial correlation approach. Using the IAB employment sample, an earlier version of the data set we use for our empirical analysis, he treats the immigration of ethnic Germans as a natural experiment and finds no detrimental impact on wages but strong displacement effects. Although the partial correlation approach in that paper differs from the equilibrium framework used here, his findings indicate that wage rigidities are prevalent in Germany.

In two other recent contributions, D’Amuri et al. (2008) and Felbermayr et al. (2008), apply a general equilibrium framework to the analysis of the labor market effects of immigration in Germany. Although both papers highlight the importance of unemployment and wage rigidities, their approach differs from ours. Similarly to Borjas (2003) and Ottaviano/Peri (2006), both papers derive the wage effects from a nested production function framework. Their identification strategy relies implicitly on the assumption that labor markets clear. Following the standard procedure in the literature, D’Amuri et al. (2008) estimate the employment effects separately from the wage effects in reduced-form equations, while our paper determines the wage and employment effects simultaneously in an equilibrium framework with imperfect labor markets. There are further important differences between these papers and ours in the identification strategy and data sources. Felbermayr et al. (2008) base their analysis on the German Socio-Economic Panel (GSOEP), which suffers from attenuation bias since the foreigner shares in specific education and experience cells are very small there. D’Amuri et al. (2008) use an earlier version of the IABS but impute the substantial immigration of ethnic Germans during the 1990s, while we identify them directly using information on their participation in labor market programs. We imputed the education variable using a procedure especially developed for the IABS that improves its quality considerably. This point is particularly relevant in case of foreign workers. Furthermore, D’Amuri et al. (2008) only estimate the elasticity of substitution between natives and foreigners and take the elasticities of substitution across education and experience cells from the US literature, while we estimate the full set of elasticities.

The remainder of this paper is organized as follows. Section 2 outlines our theoretical framework. Section 3 describes the data set. Section 4 presents the identification strategy and the estimation results for the elasticities of the wage-setting curves and the parameters of the production function. Section 5 simulates the employment and wage impact of immigration on the German labor market. Finally, Section 6 concludes.

II. Theoretical background

Building on Boeri/Brücker (2005) and Levine (1999) we apply a wage-setting framework to analyze the wage and employment effects of immigration (Layard/Nickell 1986, Layard et al. 2005). Our model replaces the conventional labor supply curve with a wage-setting function, which is consistent with right-to-manage models of collective bargaining (Nickell/Andrews 1983) and efficiency wage theories derived from turnover cost (Salop 1979) or shirking (Shapiro/Stiglitz 1984) models. We do not present an explicit collective bargaining or efficiency wage model here, since both types of models are relevant in the German context. Depending on the bargaining power of workers or human capital characteristics relevant to efficiency wage considerations, the slope of the wage-setting curve may vary in the different segments of the labor market. We therefore follow a suggestion by Card (1995) and allow the elasticity of the wage-setting curve to differ across education and experience groups of the labor market.

The wage-setting model

Suppose an economy where identical firms produce varieties of a differentiated good under monopolistic competition using different types of labor, $L_\ell,\ \ell = 1, \dots, n$, and physical capital. Production involves some fixed setup costs, but thereafter each firm produces output with constant returns to scale. Hence, production of a representative firm is given by

$$Y^i = F\left(\mathbf{L}^i, K^i\right), \tag{1}$$

where Y^i denotes a variety of the output good, $\mathbf{L}^i$ the vector of labor inputs, K^i physical capital and the superscript i the firm index. The production technology $F(\cdot)$ is increasing, concave, twice continuously differentiable in all inputs and homogeneous of degree one.

Let $\bar{N}_\ell, \ell = 1, \dots, n$ be the pre-migration labor force in each cell of the labor market. The post-migration labor force is then given by

$$N_\ell = \bar{N}_\ell + \gamma_\ell M, \qquad \textstyle\sum_\ell \gamma_\ell \ = \ 1, \tag{2}$$

where M is the total labor influx and γ_ℓ is the share of workers of type ℓ in the total immigrant inflow.

Wages and the demand for labor are determined sequentially. In the first stage, wages are fixed, whereas in the second stage, given the agreed wages, firms set prices and hire workers up to a level where profits are maximized. Suppose that each firm faces a constant elasticity of demand $\eta > 1$. Profit maximization implies that the wage rate equals

$$w_\ell^i = \nu^{-1} P^i Y_{L_\ell}^i, \quad \forall \ell,$$

where the mark-up, ν, is $(1 - 1/\eta)^{-1}$, P^i the product price of variety i of the output good, and $Y_{L_\ell}^i$ the marginal product of labor.

With identical firms we can move to the level of the aggregate economy by writing $w_\ell^i = w_\ell$, $Y_{L_\ell}^i$=Y_{L_ℓ}, and $P^i = P = 1$, where we have normalized the price level to one. The real wage is then given by

$$w_\ell = \nu^{-1} Y_{L_\ell}, \quad \forall \ell. \tag{3}$$

Equation (3) determines the demand for labor given the real wage. In the first stage of the decision process, firms and employees set wages such that they decline if the unemployment rate increases. This enables us to write the aggregate wage-setting equation as

$$w_\ell = \phi_\ell(u_\ell), \quad \phi_\ell' < 0, \quad \forall \ell, \tag{4}$$

where ϕ_ℓ is a function that captures the response of the wage to the unemployment rate $u_\ell = 1 - L_\ell / N_\ell$. The rationale behind equation (4) is that a higher unemployment rate weakens the outside options of workers and, hence, reduces their wages. As outlined above, this is consistent with a range of wage-setting models including right to manage collective bargaining models and efficiency wage models.

The wage-setting relation in equation (4) and the relation between the real wage and marginal product of labor in equation (3) allow us to solve for the employment response to a change in foreign labor supply. This requires solving a system of equations which is determined by the wage-setting curves and the production function for each type of labor. This system has to satisfy, in each cell of the labor market, the implicit function

$$\Omega_\ell(\mathbf{L}, M) \equiv \nu^{-1} Y_{L_\ell}\Big(\mathbf{L}, K\big(\mathbf{N}(M)\big)\Big) - \phi_\ell\Big(u_\ell\big(L_\ell, N_\ell(M)\big)\Big) = 0, \quad \forall \ell. \tag{5}$$

Note that equation (5) implies that the capital stock may adjust to labor supply shocks, i.e., that $\partial K/\partial \mathbf{N} \geq 0$. Differentiating this system implicitly with respect to a marginal labor supply shock through immigration yields for the change in employment

$$\frac{d\mathbf{L}}{dM} = \left(\frac{\partial \nu^{-1}\mathbf{Y_L}}{\partial \mathbf{L}} - \frac{\partial \boldsymbol{\phi}}{\partial \mathbf{u}}\frac{\partial \mathbf{u}}{\partial \mathbf{L}}\right)^{-1} \times \left(\frac{\partial \boldsymbol{\phi}}{\partial \mathbf{u}}\frac{\partial \mathbf{u}}{\partial \mathbf{N}}\frac{d\mathbf{N}}{dM} - \frac{\partial \nu^{-1}\mathbf{Y_L}}{\partial K}\frac{\partial K}{\partial \mathbf{N}}\frac{d\mathbf{N}}{dM}\right), \tag{6}$$

where $\mathbf{Y_L}$ denotes a vector of the marginal products of labor in each cell of the labor market as outlined in equation (3), $\boldsymbol{\phi}$ a vector of functions that determines the wage response to the unemployment rate as outlined in equation (4), and $\mathbf{u}$ the vector of unemployment rates.

Having solved for the equilibrium employment response, it is straightforward to use the relation in equation (3) for deriving the wage effects of migration:

$$\frac{d\mathbf{w}}{dM} = \frac{\partial \nu^{-1}\mathbf{Y_L}}{\partial \mathbf{L}}\frac{d\mathbf{L}}{dM} + \frac{\partial \nu^{-1}\mathbf{Y_L}}{\partial K}\frac{\partial K}{\partial \mathbf{N}}\frac{d\mathbf{N}}{dM}. \tag{7}$$

Outline of the empirical framework

For the empirical analysis, we have to impose more structure on the economy. Similar to Borjas (2003) and Ottaviano/Peri (2006), we follow Card/Lemieux (2001) in using a nested CES production function. More specifically, we employ a four-level production function which groups the workforce into $q = 1, \ldots, 4$ education groups, $j = 1, \ldots, 8$ experience groups, and $k = 1,2$ nationality groups.[1] Although the nested CES function imposes some restrictions on the elasticities of substitution, it has the advantage that it is parsimonious in the parameters.

Suppose that aggregate production in equation (1) can be represented by a standard Cobb-Douglas production function:

$$Y_t = A_t L_t^{\alpha} K_t^{1-\alpha}, \tag{8}$$

where Y_t denotes aggregate output, A_t an exogenous parameter which captures total factor productivity, L_t the aggregate labor input, K_t physical capital, α the production elasticity of la-

[1] Our four-level framework resembles the specification by Ottaviano/Peri (2006). D'Amuri et al. (2008) have recently applied a five-level framework to Germany that distinguishes between different vintages of immigrants in a further nest of the production function. They find that old and new arrivals are perfect substitutes.

bor, and t the time index.

Grouping the labor force by education, experience, and national origin yields:

$$L_t = \left[\sum_{q=1}^{4} \theta_{qt} L_{qt}^{(\delta-1)/\delta}\right]^{\delta/(\delta-1)}, \qquad \sum_{q=1}^{4} \theta_{qt} = 1, \tag{9}$$

$$L_{qt} = \left[\sum_{j=1}^{8} \theta_{qj} L_{qjt}^{(\rho-1)/\rho}\right]^{\rho/(\rho-1)}, \qquad \sum_{j=1}^{8} \theta_{qj} = 1, \tag{10}$$

$$L_{qjt} = \left[\sum_{k=1}^{2} \theta_{qjk} L_{qjkt}^{(\sigma_q-1)/\sigma_q}\right]^{\sigma_q/(\sigma_q-1)}, \sum_{k=1}^{2} \theta_{qjk} = 1, \tag{11}$$

where the aggregate L_t incorporates the contributions of workers who differ in both education and experience, L_{qt} is a labor composite that aggregates all workers with education q, L_{qjt} a labor composite that aggregates native and migrant workers of education q and experience j, and L_{qjkt} the number of employed workers of education q, experience j, and national origin k. The technology parameters θ_{qt}, θ_{qj},, and θ_{qjk} determine the productivity levels of the respective factor. We allow the productivity parameter θ_{qt} to vary over time since skill-biased technological progress might affect the productivity of various types of labor in different ways (Katz/Murphy 1992). The other production parameters are assumed to be constant over time.

Finally, $\delta > 0$, $\rho > 0$, and $\sigma_q > 0$ are constant parameters measuring the elasticity of substitution between labor of different educational levels, between workers with similar education but different work experience, and between native and migrant workers with similar education and experience levels. Our a priori expectation is that workers within each experience group are closer substitutes than those across skill groups, which implies that $\rho > \delta$.

Whether foreign and native workers in each education and experience group are imperfect substitutes is the subject of some controversy in the literature. We therefore test empirically whether native and foreign workers are imperfect substitutes.

Based on equation (3) we can write the real wage rate as the marginal product of labor divided by the mark-up factor. Using the nested CES production function we thus write the log wage of a worker with skill q, experience j and national origin k as

$$
\begin{aligned}
\ln w_{qjkt} &= \ln\left(\nu^{-1}\alpha A_t^{1/\alpha}\right) + \frac{1}{\delta}\ln L_t + \ln\Theta_{qt} - \left(\frac{1}{\delta}-\frac{1}{\rho}\right)\ln L_{qt} + \ln\Theta_{qj} \\
&\quad - \left(\frac{1}{\rho}-\frac{1}{\sigma_q}\right)\ln L_{qjt} + \ln\Theta_{qjk} - \frac{1}{\sigma_q}\ln L_{qjkt} + \frac{1-\alpha}{\alpha}\ln\kappa_t
\end{aligned}
\tag{12}
$$

where $\kappa_t \equiv K_t/Y_t$ denotes the capital-output ratio.

To calculate the wage effects of a labor supply shock due to immigration, we first compute the employment effects. The general solution for the employment effects is given in equation (6), and an explicit solution for our case with 64 types of labor and a nested CES production function is provided in Appendix A1. In the second step, we differentiate the wage equation (12) with respect to the employment changes in all cells of the labor market and with respect to a change in the capital-output ratio triggered by immigration. This gives

$$
\begin{aligned}
\frac{dw_{qjkt}}{w_{qjkt}} &= \frac{1}{\delta}\sum_z\sum_x\sum_m\left(s_{zxmt}\,\frac{dL_{zxmt}}{L_{zxmt}}\right) - \left(\frac{1}{\delta}-\frac{1}{\rho}\right)\frac{1}{s_{qt}}\sum_x\sum_m\left(s_{qxmt}\,\frac{dL_{qxmt}}{L_{qxmt}}\right) \\
&\quad - \left(\frac{1}{\rho}-\frac{1}{\sigma_q}\right)\frac{1}{s_{qjt}}\sum_m\left(s_{qjmt}\,\frac{dL_{qjmt}}{L_{qjmt}}\right) - \frac{1}{\sigma_q}\left(\frac{dL_{qjkt}}{L_{qjkt}}\right) + \frac{1-\alpha}{\alpha}\frac{d\kappa_t}{\kappa_t},
\end{aligned}
\tag{13}
$$

where $z = 1, \dots, 4$ indexes education, $x = 1, \dots, 8$ work experience, and $m = 1,2$ national origin, and s denotes the share of wage sum paid to workers in the respective labor market cell in the total wage bill, i.e.,

$$
s_{qjkt} = \frac{w_{qjkt}\,L_{qjkt}}{\sum_z\sum_x\sum_m w_{zxmt}\,L_{zxmt}},
$$

$$
s_{qjt} = \frac{\sum_m w_{qjmt}\;L_{qjmt}}{\sum_z\sum_x\sum_m w_{zxmt}\,L_{zxmt}} \quad \text{and}
$$

$$
s_{qt} = \frac{\sum_x\sum_m w_{qxmt}\;L_{qxmt}}{\sum_z\sum_x\sum_m w_{zxmt}\,L_{zxmt}}.
$$

III. Data and descriptive evidence

In our empirical analysis we use the IAB employment sample (IABS), a two percent random sample of all German employees registered with the social security system during the period

1975-2004. In addition to socio-economic and job characteristics, the IABS provides information on benefit recipients at the individual level.[2] The IABS is stratified according to nationality and therefore representative of the native and foreign working population. The data set is especially useful for analyses that take wages into account since the wage information is used to calculate social security contributions and is therefore highly reliable.[3]

Nevertheless the IABS also has some limitations in the context of our analysis, the main one being that we can identify foreigners only on the basis of citizenship. Some further shortcomings arise from the wage and qualification information provided by the data set.

First, there is no information on the year when immigrants entered the country. Due to the *jus sanguinis* tradition of the German law, naturalization rates are traditionally very low, such that second- and third-generation migrants often still have foreign citizenship and therefore appear as foreign workers in our sample. On August 1, 1999, a new immigration act came into effect that allows German-born children of foreign-born parents who have been living in Germany for at least eight years to decide which nationality to adopt up to the age of 23. This has slightly increased the naturalization of German-born individuals with a migrant background. To mitigate the possible effects of naturalization, we have classified all individuals as foreigners who are reported as foreign citizens in their first available spell. This prevents naturalizations from being displayed as a declining foreigner share in our sample.

Second, ethnic Germans – so-called "Spätaussiedler" – are reported as Germans since the concept of citizenship does not allow us to distinguish between citizens born in Germany and those born abroad. However, language courses and other integration subsidies offered to ethnic Germans should facilitate their labor market integration. These programs are reported in the benefit recipient file added to our data set. This allows us to identify the majority of ethnic

2 About 90 percent of the registered unemployed are eligible for benefits (Wagner/Jahn 2004). Therefore the unemployment rate is only slightly downward-biased.

3 In our sample the average size of the foreigner cells is well above 1,000 observations. Therefore it is not likely that our results suffer from an "attenuation bias" (Aydemir/Borjas 2006).

Germans who have entered the German labor force since 1980. In our sample, the cumulative inflow of ethnic Germans accounts for 3.2 percent of the labor force in Western Germany.[4] Since ethnic Germans' labor market performance and language command resembles that of other foreigners (see Glitz 2006), we have classified ethnic Germans as foreigners.

Third, the IABS included Eastern Germany for the first time in 1992. Solely focusing on the unified Germany would exclude the main immigration shock associated with the fall of the iron curtain. German reunification also requires excluding Western Berlin, since mobility between Eastern and Western Berlin has been high since the fall of the Berlin wall in 1989. We therefore concentrate in our analysis to individuals who were employed or unemployed in West Germany on September 30 of any year in the period 1980 to 2004. We do not believe that the focus on Western Germany should significantly affect our results as four-fifth of the German labor force work in Western Germany and the foreigner share is negligible in Eastern Germany.

Fourth, the data set reports gross daily wages and does not provide information on hours worked. We therefore exclude part-time employees, trainees, interns, and at-home workers from the sample since the wage information is not comparable for these groups. For the same reason we exclude workers with wages below the social security contribution threshold.

Fifth, there is some empirical evidence of differences in the early retirement behavior between German and immigrant men (Bonin et al. 2000). We therefore restrict our analysis to individuals between the ages of 15 and 60.

Sixth, our data are right-censored since gross wages can only be observed up to the social security contribution ceiling. About four percent of the employment spells in the final data set are right-censored. This may affect the estimation of the wage-setting curves, particularly in the high-skilled segments of the labor market. We have therefore imputed wages above the social security contribution ceiling using a heteroscedastic single imputation approach specifically de-

[4] Appendix B1, which is available upon request, provides information on the stock and flow of ethnic German workers into the German labor market.

veloped for the IABS data set (Büttner/Rässler 2008).[5]

Seventh, self-employed workers and civil servants do not contribute to the social security system and are therefore not covered by our sample. While the self-employment rate of natives increased only slightly from 9.5 percent in 1985 up to 10 percent in 2000, the self-employment rate of immigrants increased from about 6 percent to about 9 percent (Kontos 2003). Nevertheless, the change has been moderate and we do not expect that this will bias our results considerably. In the case of civil servants, it seems plausible to assume that due to legal restrictions, immigrants do not displace natives.

Eighth, the information on education is provided by employers. This means that information on educational levels is missing for about 17 percent of the individuals. Foreigners are disproportionately affected by missing information on educational levels. We therefore imputed the missing information on education by employing a procedure developed by Fitzenberger et al. (2005), which allows inconsistent education information to be corrected over time as well. After applying this imputation procedure, we had to drop only 1.6 percent of the individuals due to missing or inconsistent information on education (see Appendix A2).

Finally, education and work experience acquired in foreign countries may not have the same value in the labor market as education and experience obtained in Germany. Moreover, certain characteristics of foreigners, such as their command of the German language, may prevent them from fully transferring their human capital to the German labor market. However, the classification of educational levels in the IABS reflects the official recognition of educational degrees acquired abroad, which is rather restrictive in Germany. As a consequence, the correlation between educational degree and occupational status is similar for foreigners and natives in our data set (in fact, slightly higher for foreigners). In the highest education group, i.e., individuals with a university degree, the share of immigrants in high-level occupations is only slightly be-

5 A more detailed description of the imputation procedure is provided in Appendix B2, which is available upon request.

low that of natives (see Glitz 2006 for a similar result). Note also that in their analysis of wage elasticities D'Amuri et al. (2008) find no differences, whether classifying the labor force by educational or by occupational level.

As the above sample selection mainly affects groups in the labor market where foreigners are underrepresented (civil servants, the self-employed, part-time workers), we might slightly overstate the impact of legal immigration in our analysis. However, one caveat might be that we are not able to observe illegal immigrant workers, who may exert pressures on the less-skilled segments of the labor market in particular.

Following the model outlined in Section 2, we group the labor force by education and potential work experience. A sensible classification following the characteristics of the German labor market requires us to distinguish four educational groups: no vocational degree, vocational degree, high school degree ("Abitur") with vocational degree, and university degree. At first glance, one might consider aggregating the groups "vocational degree" and "high school degree with vocational degree", but in Germany these are separate labor markets. Despite the small size of the group with a high school degree, we therefore decided to treat it separately.

Furthermore, we distinguish eight potential work experience classes following the standard approach by Borjas (2003), subtracting the typical number of years spent in the educational system from the age of the worker and splitting experience into five-year intervals.

Figure 1 displays the share of foreigners – including ethnic Germans – in the labor force and the share of foreigners in the employed workforce. During the 1980s we observe a sharp decline, which is a consequence of tightening migration restrictions in the wake of the first and second oil price shock. Between the mid 1980s and the mid 1990s, the foreign share in the labor force increased by 5 percentage points. The sharp increase in the foreigner share during the 1990s resulted from the collapse of the Central and Eastern European communist regimes and the civil wars in the former Yugoslavia, which triggered large-scale migration to Germany. Since the early 2000s, the foreigner share has plateaued due to the slowdown in economic growth and tighter restrictions on immigration. Moreover, since foreigners tend to be more than proportionally affected by unemployment, their share in the employed workforce declined relative to their share in the labor force at the end of the millennium.

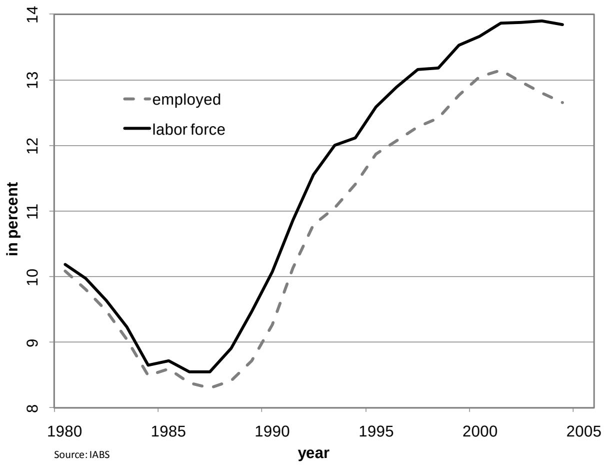

Figure 1: Share of foreign work force

Table 1 shows that the increase in the foreign labor supply disproportionally affects the higher education groups: while the number of foreign workers without a vocational degree declined by one-fourth during the 1980-2004 period, the number of foreigners with a university degree increased by 164 percent, and the number of foreigners with a high-school and vocational training degree by a factor of 6. The largest education group – those with a vocational degree – increased by 134 percent, while the total foreign labor force increased by 40 percent.

[Table 1 about here]

IV. Estimation

Wage-setting equations

The first step in the empirical application of the model outlined in Section 2 is to estimate the wage-setting equations. Following Bell et al. (2002) and Blanchflower/Oswald (2005) we estimate the elasticity of the wage-setting curve in dynamic form, i.e., as

$$\ln w_{qjt} = \beta_{qj} \ln w_{qj,t-1} + \gamma_{qj} \ln u_{qjt} + \lambda_{qj} \tau_{qjt} + \boldsymbol{\eta}' \mathbf{X}_t + e_{qjt}, \qquad (14)$$

where γ_{qj} denotes the short-run elasticity between the wage and the unemployment rate, $\mathbf{X}_t$ a vector of control variables, $\boldsymbol{\eta}$ the corresponding vector of coefficients, and τ_{qjt} an education-experience-specific deterministic time trend. We include a linear and a squared trend here. As controls for macroeconomic shocks we use real GDP, the consumer price index, and the industrial production index.[6] The error term e_{qjt} is specified as a one-way error component model with fixed effects for each education-experience group.

Unobserved shocks may affect wages and the unemployment rate simultaneously. In order to address this problem, we estimate equation (14) by 2SLS. We use three instruments for the unemployment rate. First, following Bartik (1991) and Blanchard/Katz (1992) we employ an industry mix variable that measures how much of the deviation in employment growth in an education-experience cell from average employment growth can be explained by a concentration of workers in the respective cell in fast- or slow-growing industries. This variable simply measures how much of the change in employment can be attributed to an exogenous shift of the sectoral structure. Our second instrument is an export demand index, which is constructed as the log of the GDP per capita at constant prices and exchange rates of all OECD countries weighted by their average share in German exports during the sample period. This variable should capture exogenous shifts in labor demand that are triggered by the economic activity of Germany's trading partners.[7]

Third, we instrument the unemployment rate with a potential immigration variable, which measures potential exogenous labor supply shocks due to immigration in each education-experience cell of the labor market. Since immigration itself might be endogenous, we estimate auxiliary regressions that explain the stock of foreign workers in each education-experience cell by push factors in the sending countries (log GDP per capita, log unemployment rate, institutional variables that capture political shocks and migration conditions) and bilateral fixed effects. The coefficients of these regressions are used to calculate the migration potential in each education-experience group. By construction, this variable captures labor supply shocks due to

6 See Appendix A3 for a description of the variables.

7 A similar instrument has been used by Carlsson et al. (2008) in the estimation of an employment equation.

immigration that are driven by exogenous factors. Several diagnostic tests support the hypothesis that our instruments are valid and relevant (see notes in Table 2).

The specification of equation (14) is similar to that used in the wage-setting and wage curve literature, but it differs from the usual approach in that it allows the elasticity between wages and the unemployment rate to differ across education-experience groups. However, we aggregate native and foreign workers.[8] This reflects not only the difficulties in empirically identifying the elasticities of the wage-setting curve for a rather small group like foreigners, but also the fact that collective wage agreements determine a wage floor for education and experience groups but do not discriminate a priori between native and foreign workers. Moreover, the German anti-discrimination legislation reduces opportunities to set wages differently for native and foreign workers. [9]

For the identification of the coefficients in equation (14) we exploit the long time dimension of our data set. The underlying assumption that the coefficients are stable over time may be questionable if the wage-setting mechanism changes. In the case of Western Germany, the number of employees covered by industry- or firm-level collective agreements did indeed decline by 11 percentage points between 1996 and 2004. However, most firms that left the employer federations still apply wage levels fixed in collective agreements, such that the share of employees who are directly or indirectly covered by collective wage agreements remained rather stable at about 85 percent of the workforce in Western Germany (Ellguth/Kohaut 2007). [10]

[Table 2 about here]

The results are displayed in Table 2. All regressions have the expected negative sign for the coefficient on the unemployment rate. All short-term elasticities are significant and with one

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8 In each regression we have pooled two experience groups to achieve more stable results.

9 We can hardly test for this assumption since we estimate point elasticities and since the unemployment rate differs between natives and foreigners.

10 We conducted Wald tests to test for potential structural breaks that might be caused by changes in the wage-setting mechanism at the middle of the sample period. The test statistics cannot reject the null hypothesis of no structural breaks.

exception, the long-term elasticities are also significant. The autoregressive parameter on the lagged wage is well below 1, supporting a wage-setting curve rather than a Phillips curve.[11]

The first regressions provide pooled estimates of the wage-setting curve for all groups and for each education group separately. In the regression where all education-experience groups are pooled, we find a short-run elasticity of about -0.08 and a long-run elasticity of about -0.16. The national-level estimates presented here are somewhat higher than the average elasticity of -0.1 found by the regional-level wage curve literature in other OECD countries (Blanchflower/ Oswald 1994, Nijkamp/Poot 2005), but much higher than the elasticity of -0.03 estimated by Baltagi et al. (2009) at the regional level in Germany. This is not surprising since the regional level estimates control for all macroeconomic influences that are particularly relevant in economies such as Germany, where industry-level bargaining plays an important role.

The most intriguing finding in our disaggregated estimates is that of very high elasticities of the wage-setting curve in the segments with little work experience, with the notable exception of the group without vocational training. This indicates that seniority wage considerations play an important role in wage-setting in the skilled segments of the labor market. Note that this finding is relevant in our context since most newly arrived immigrants possess little work experience. Interestingly enough, we do not obtain a lower elasticity of the wage-setting curve in the group where union density and the coverage rate of collective wage agreements is particularly high in Germany: individuals with vocational training. In contrast, the aggregate elasticities of the wage-setting curve are, at values of between -0.15 and -0.19, relatively similar across educational groups.

Labor demand equations

In the next step we estimate the elasticities of substitution between the different types of labor based on our nested production function framework. Our identification strategy differs from the

[11] In one case we obtain a negative but insignificant coefficient for the lagged wage. In our simulations we assume that adjustment takes place immediately, in this case by setting this coefficient to zero.

one used in the literature, which relies on perfect competition with clearing labor markets (Card/Lemieux 2001, Borjas 2003, Ottaviano/Peri 2006, 2008). While this literature treats employment as exogenous and wages as the endogenous variable, it follows from our wage-setting framework that labor demand is endogenously determined once wages are fixed.

Let us start with the identification of the elasticity of substitution between native and foreign workers. Based on equation (12) we can express the relative demand for native and foreign workers with education q and experience j as $\ln(L_{qjht}/L_{qjft}) = \sigma_q \ln(\theta_{qjh}/\theta_{qjf}) - \sigma_q \ln(w_{qjht}/w_{qjft})$, where the index h denotes natives and the index f foreigners. We estimate this equation as

$$\ln(L_{qjht}/L_{qjft}) = D_{qj} - \sigma_q \ln(w_{qjht}/w_{qjft}) + \boldsymbol{\mu}'\mathbf{X}_t + \varepsilon_{qjt} \qquad (15)$$

where D_{qj} denotes a vector of dummy variable for each education-experience cell, $\mathbf{X}_t$ a vector of control variables, $\boldsymbol{\mu}$ the corresponding vector of coefficients and ε_{qjt} a zero-mean disturbance term. Following the approach of Ottaviano/Peri (2008) the dummy variables in each education-experience cell capture the log of the relative labor productivity of natives and foreigners times the elasticity of substitution. This implies that the relative productivity of natives and immigrants varies across education and experience groups but is constant over time.[12]

As macroeconomic control variables we use real GDP growth, an index of the domestic crude oil price, and the export performance index of the OECD. We do not consider time fixed effects since the model is identified by the variance over time. Note that technology shifts in the productivity of education (or experience) groups and macroeconomic shocks common to both natives and foreigners are absorbed by higher levels of the production function.

Estimating equation (15) by OLS can generate inconsistent results if unobserved idiosyncratic shocks affect both the relative labor demands and relative wages of natives and foreigners. To

[12] Borjas et al. (2008) suggest also including interaction dummies of the education-experience-specific fixed effects with linear time trends, which would absorb a large part of the identifying variation. However, there is no empirical evidence for Germany that the relative labor productivity of foreigners has changed systematically over time at a given level of education and work experience.

address this problem we estimate equation (15) with 2SLS. As instruments we first use the ratio of the log of the average level of unemployment benefits for natives and foreigners in each education-experience cell as an instrument for the relative wage levels. Accurate information on the level of unemployment benefits is provided by the IABS. Unemployment benefits are a suitable instrument if they affect the wage level via the reservation wage without moving labor demand directly. Since it may take time for unemployment benefits to affect wages, we use the first and the second lag of the unemployment benefits as instruments.

Our second instrument is an ideology index, which captures the share of left- and right-wing parties in the government weighted by their seats in parliament (Bjørnskov 2008)[13]. This instrument captures governmental policies that can affect reservation wages through different channels, e.g., progressive taxation, generosity of welfare benefits, etc. Note that foreigners' access to welfare benefits is a core issue in the policy debate, which in turn affects relative reservation wages for foreigners and natives..[14]

[Table 3 about here]

The results of our estimates are reported in Table 3. All coefficients for σ_q are significantly different from zero, providing support for the hypothesis that native and foreign workers are imperfect substitutes. At 7, the overall elasticity between native and foreign workers is similar to that found by Ottaviano/Peri (2006) for the US, but smaller than the elasticity of between 16 and 21 identified by D'Amuri et al. (2008) for Germany based on an identification strategy that assumes clearing labor markets. The elasticity of substitution is particularly high for workers with vocational training and a high school degree, but relatively low both for less skilled workers and workers with a university degree.

In the next step we estimate the elasticity of substitution between experience groups. Using

[13] We are grateful to Christian Bjørnskov who provided the ideology index.

[14] The regression diagnostics indicate that these instruments are both valid and relevant (see notes to Table 3). In addition, we conducted Wald tests to test for potential structural breaks at the middle of the sample period. The test statistics cannot reject the null hypothesis of no structural breaks.

equation (12) once again we can estimate the demand for the labor composite L_{qjt} as

$$\ln \hat{L}_{qjt} = D_t + D_{qt} + D_{qj} - \rho \ln w_{qjt} + \upsilon_{qjt}, \tag{16}$$

where the time-specific fixed effects D_t control for the variance of $\rho \ln\left(\nu^{-1}\alpha A_t^{1/\alpha}\kappa_t^{(1-\alpha)/\alpha}\right) + \rho/\delta \ln L_t$, the time by education-specific fixed effects D_{qt} for the variation in $\rho \ln \Theta_{qt} - (\rho/\delta - 1)\ln L_{qt}$, and the education-experience group fixed effects D_{qj} for the variance in $\rho \ln \Theta_{qj}$, i.e., in the productivity term times the elasticity of substitution, which is assumed to be constant over time. υ_{qjt} denotes the zero-mean disturbance. The labor composite is calculated as $\hat{L}_{qjt} = \left[\hat{\Theta}_{qjh} L_{qjht}^{(\hat{\sigma}_q-1)/\hat{\sigma}_q} + \hat{\Theta}_{qjf} L_{qjft}^{(\hat{\sigma}_q-1)/\hat{\sigma}_q}\right]^{\hat{\sigma}_q/(\hat{\sigma}_q-1)}$, where we use our estimates of the education-experience-specific fixed effects from equation (15) to calculate the productivity parameters for native and foreign workers as $\hat{\Theta}_{qjh} = \frac{\exp(\hat{D}_{qj}/\hat{\sigma}_q)}{1+\exp(\hat{D}_{qj}/\hat{\sigma}_q)}$ and $\hat{\Theta}_{qjf} = \frac{1}{1+\exp(\hat{D}_{qj}/\hat{\sigma}_q)}$.

We estimate equation (16) by 2SLS using the first and second lags of the log of the average unemployment benefit in each education-experience cell and the ideology index as instruments. Our regression diagnostics again suggest that these instruments are valid and relevant. We find an elasticity of substitution across experience groups of 8.6, which is close to other findings in the international literature (Card/Lemieux 2001, Borjas 2003, Ottaviano/Peri 2006).

The elasticity of substitution between education groups is estimated analogously as

$$\ln \hat{L}_{qt} = D_t + D_q + \beta_q \tau_{qt} - \delta \ln w_{qt} + \xi_{qt}, \tag{17}$$

where the time-specific fixed effects D_t control for the variance of $\delta \ln\left(\nu^{-1}\alpha A_t^{1/\alpha}\kappa_t^{(1-\alpha)/\alpha}\right) + \ln L_t$ and other macroeconomic fluctuations, the education-specific fixed effects D_q, and the education-specific deterministic time trend τ_{qt} for the variance in the term $\delta \ln \Theta_{qt}$, which captures the variance in the skill-specific productivity parameter which is driven inter alia by skill-biased technological progress (see Katz/Murphy 1992, for a similar specification). ξ_{qt} denotes a zero-mean random disturbance. The labor composite $\hat{L}_{qt}$ is computed as

$\hat{L}_{qt} = \left[\sum_{j=1}^{8} \hat{\theta}_{qj} \hat{L}_{qjt}^{(\hat{\rho}-1)/\hat{\rho}}\right]^{\hat{\rho}/(\hat{\rho}-1)}$, where the estimated efficiency parameters $\hat{\theta}_{qj}$ are derived from the fixed-effects estimates as $\hat{\theta}_{qj} = \frac{\exp(\hat{D}_{qj}/\hat{\rho})}{\sum_j \exp(\hat{D}_{qj}/\hat{\rho})}$.

We estimate equation (17) by 2SLS using the log of first and second lags of the average unemployment benefits in each education group as an instrument, which again turns out to be valid and relevant. We find an elasticity of 3, which is similar to that found in the US literature (Katz/Murphy 1992, Ottaviano/Peri 2006), but below what other studies find for Germany (Felbermayr et al. 2008).

V. Simulation of the immigration impact

We now use the estimated parameter values to simulate the impact of migration on (un)employment and wages. We simulate two scenarios here.

First, we simulate the effects of a one percent increase in the labor force due to immigration using the education and experience composition of the foreign workforce at the average of the sample period. This scenario provides an indication as to the marginal effects of immigration at the given structure of the foreign workforce.

Second, we simulate the annual averages of the wage and employment effects of the actual immigration shock for the entire 1980-2004 period. During this period the foreign labor force increased by 40 percent, which corresponds to an increase of four percent of the total labor force. Since this is an inframarginal labor supply shock in many cells of the labor market, we calculate annual averages for this period. For this purpose we first simulate the annual wage and unemployment effects based on the actual changes in each cell of the labor market for each year and then calculate the annual average for the 1980-2004 period using the share of the wage sum and of the labor force in each education and experience cell in each year as weights.[15]

[15] Appendix B3, which is available upon request, also provides the results dividing the observation period in five year intervals. The results do not differ fundamentally, except for the subperiods 1985-1989 and 1990-1994, where the immigration shock has been particularly large. For these two intervals we find that the unemployment rate increases and wages drop in all foreign education groups.

In the one percent scenario, we distinguish between the short-run and the long-run effects of immigration. In the simulations of the short-run impact we use the short-term elasticities of the wage-setting curves as estimated in equation (14) and assume that the capital stock remains fixed. In the long-run simulations we use the long-term elasticities of the wage-setting curves and assume that the capital stock adjusts completely to an aggregate labor supply shock, such that the capital-output ratio is fixed. The latter assumption is empirically supported by the Kaldor facts on economic growth and can be explained, inter alia, by international capital mobility (see Ottaviano/Peri 2006, for a discussion). Finally, as a robustness check, we have calculated the effects of immigration assuming clearing labor markets, i.e. an economy where the elasticity of the wage-setting curve approaches infinity and unemployment is zero.

In the scenario that covers the 1980-2004 period we present the long-run scenario only, since a short-term scenario does not make sense for such a long time period.

In all scenarios, we start with the calculation of the change in employment based on equation (6). The explicit form of the vectors and matrices that we employ on the basis of our nested production function is provided in Appendix A1. The parameters for the wage-setting curves are taken from our estimates of equation (14) and the parameters σ_q, ρ, and δ from our estimates of equations (15) – (17). Following the literature, we set α to 0.67 (Cahuc/Zylberberg 2001). Having calculated the employment effects of immigration, we use equation (13) for the calculation of the wage effect. The shares of education and education-experience groups in the total wage bill are taken from our data set. Note that the mark-up factor is a constant that cancels out when we calculate changes of wage and (un)employment levels.

Table 4 reports the average effects for the total labor force, the native labor force, and the foreign labor force by educational levels. For the calculation of the average effects, we weight the wage changes by the income share in each cell, and the changes in the unemployment rate by the share in the labor force in each cell.

Effects of a one percent increase: Our simulation results indicate that an immigration of one percent of the labor force at the average skill and experience structure of the foreign workforce during the sample period reduces overall average wages by 0.18 percent and increases the average unemployment rate by 0.31 percentage points in the short run. The pre-existing foreign labor force bears the brunt of adjustment: their wages decline by 1.11 percent and their unemployment rate increases substantially by almost two percentage points in the short run. In contrast, the native workforce is only slightly affected. Their wages decline by 0.08 percent and their unemployment increases by 0.09 percent in the short run.

The large difference in the labor market effects for the foreign and the native workforce can be traced back to two main facts: First, the elasticity of substitution between native and foreign workers is relatively low, which implies that the labor supply shocks in specific education and experience cells of the labor market can spill over to native workers only to a limited extent. The particularly low elasticity of substitution between foreigners and natives in the education groups without vocational training and with a university degree is also responsible for the fact that the foreign labor supply shock mainly affects the foreign workforce in these two groups.

Second, the skill and experience structure of the immigrant influx resembles that of the foreign workforce, but differs substantially from that of the native workforce in this scenario. Since the education and experience groups are imperfect substitutes, this again implies that the foreign workforce is affected more than the native workforce. At the aggregate level, we find that the education group without vocational training suffers particularly from the labor supply shock since immigrants are represented disproportionally in this skill group.

The picture looks brighter if we consider the long-run effects of immigration. The adjustment of capital stocks ameliorates the labor supply shock, such that the average wage level of the total labor force remains constant. The unemployment rate increases by 0.08 percentage points in the long-run simulations, which can be explained by the fact that in this scenario, immigrants enter labor market cells with high unemployment rates and rather low elasticities of the wage-setting curves.

As a consequence of the low elasticity of substitution between natives and foreigners and the skill composition of the labor supply shock, the native labor force benefits in terms of higher wages (0.11 percent) and lower unemployment (-0.06 percentage points) in the long-run scenario, while foreign workers suffer from lower wages (-1.09 percent) and higher unemployment (1.16 percentage points). Again, the pre-existing foreign workforce is particularly affected in those education segments of the labor market where the elasticity of substitution between natives and foreigners is low.

Comparing these results with the counterfactual case of an economy with clearing labor markets shows that wage rigidities play an important role in protecting employed workers against wage competition from immigrants: The wage effects of immigration would increase by a factor of about 1.55 in the counterfactual case of clearing labor markets compared to the case with wage rigidities in the short-term scenario. The role of wage rigidities is mitigated, however, in the long-term scenario when capital stocks adjust: The wage effects in the clearing labor market scenario exceed those in the scenario with wage rigidities by a factor of 1.3 in the long-run according to our simulations.

[Table 4 about here]

Effects of the 1980-2004 immigration shock: In contrast to the previous scenario, the actual immigration during the 1980-2004 period is associated with a decline in the group of no vocational training, while the number of immigrants in the other skill groups increased continuously. This yields a different picture.

The most intriguing finding in this scenario is that immigration did not affect the unemployment rate in Germany. While the unemployment rate of natives and the total labor force remained constant over the period, the unemployment rate of the pre-existing foreign labor force increased only marginally – by less than 0.01 percentage points per annum. While the increasing labor supply of foreigners raised unemployment rates in the cells that were affected most by the influx, the cross-effects of the labor supply shifted the increased labor demand to other cells, such that the overall effects cancel one another out in this scenario. This is because foreign labor supply declined in the labor market segment with the highest unemployment rate and a rather low elasticity of the wage setting curve – i.e., individuals without vocational training – while increasing in the higher-skilled segments of the labor market, where unemployment is lower and the elasticity of the wage-setting curve is higher.

The wage levels of the native labor force increased slightly (+0.01 percent p.a.) while those of the foreign workforce declined slightly (-0.09 percent p.a.). We observe substantial increases in the wages of foreign workers without vocational training, while those of foreign workers with a university and a high school degree declined substantially as a consequence of the high labor influx in these skill groups. The native labor force benefits both at the lower and higher end of the skill spectrum. At the level of the entire labor force, we find that wage levels increased in the group without vocational training, and declined slightly in all other education groups, which corresponds again to the skill composition of the labor supply shift.

Comparing our findings: The elasticity between wages and immigration in our one percent scenario is about half the size of that found by Borjas (2003) and Ottaviano/Peri (2006, 2008) for the US and slightly below what Manacorda et al. (2006) and Dustmann et al. (2008) find for the UK. This is hardly surprising since these studies derive the wage effects of immigration assuming clearing labor markets, whereas we consider wage rigidities.

Comparing our findings with those of D'Amuri et al. (2008) provides some interesting insights into the differences between their outcomes and ours resulting from the different methodological approaches used to address the unemployment effects of immigration. Their reduced-form estimates of the employment impact of immigration suggest that a ten percent increase in the foreign workforce increases the unemployment rate of foreign workers by between 1.5 and two percentage points, while the native workforce remains unaffected.

This is consistent with our finding that a one percent increase of the labor force through immigration – which corresponds roughly to a ten percent increase in the foreign labor force – increases the unemployment rate of foreigners by two percentage points in the short term at the given skill structure of the foreign workforce.

However, while the partial correlation approach of D'Amuri et al. (2008) suggests that the actual migrant influx into Germany increased the unemployment rate of the foreign workforce substantially, we find that the unemployment rate of the foreign workforce remained (almost) constant over the entire sample period. This can be traced back to the fact that the increasing labor supply in the higher skilled segments of the labor market reduced unemployment in the less-skilled segments, which are characterized by lower wage flexibility. The decline in unemployment in these cells exceeded the replacement effects in the high-skilled cells of the foreign labor force. The consideration of these cross-effects in our approach thus delivers a different picture of the unemployment effects of the actual immigrant influx into Germany.

VI. Conclusions

Concerns about immigration affecting not only wages but also native employment opportunities are widespread in continental Europe, where labor market rigidities are prevalent. In this paper we present a general equilibrium framework that allows us to analyze the wage and employment effects of migration simultaneously in a setting with imperfect labor markets. Our empirical findings suggest that the wage flexibility varies widely for different segments of the labor market. The elasticity of the wage-setting curve is particularly high for workers with little work experience, i.e. labor market segments where newly arrived immigrants are more than proportionally represented.

Our approach provides a number of new insights. In contrast to the literature studying the employment effects of immigration based on partial correlations between the (un)employment and the immigrant rate in certain segments of the labor market, we find that immigration can either raise or reduce unemployment depending on the education and experience structure of the immigrant influx and the wage flexibility in different segments of the labor market. According to our simulations, the immigration of about four percent of the labor force during the 1980-

2004 period did not increase either the aggregate unemployment rate of the workforce or the unemployment rate of the foreign labor force. This can be traced back to the fact that the average skill level of the immigrant workforce has increased substantially over time. The higher labor supply in labor market segments with higher wage flexibility and lower unemployment has created an additional demand for less-skilled workers, which compensates for the replacement effects in other education and experience cells of the labor market.

Another intriguing finding is the strong evidence that native and foreign workers are imperfect substitutes in the labor market. As a consequence, the native workforce tends to benefit from immigration in terms of higher wages and lower unemployment risks in all simulations, at least in the long run, although these effects are small. In contrast, the immigration of foreign workers has a major impact on the foreign labor force. While the foreign workforce would suffer substantially from immigration at the given skill and experience structure, our results also show that the foreign workers can benefit if new immigrants are high-skilled, since immigrants are generally more than proportionally represented in the less-skilled segments of the labor market.

Our findings have some important policy implications. Selection of immigrants by human capital characteristics such as education and age is a crucial issue in economies that suffer from wage and other labor market rigidities. In the case of Germany, the gains from immigration are particularly large if immigrants are educated and if they are young, since the flexibility of the labor market is high in these segments. Moreover, policy measures that attempt to increase the elasticity of substitution between native and foreign workers, e.g., through improved labor market integration and efforts to facilitate the transfer of human capital, would mitigate the polarization of wages and employment opportunities of native and foreign workers. Although such policies would reduce the gains of the native labor force from immigration in terms of higher wages and lower unemployment, they would increase social cohesion and reduce the potential costs of immigration that arise through the welfare state channel in the receiving countries.

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Table 1: Native and foreign labor force by education group

	1980	1990	2000	2004	average 1980-2004	change 1980-2004 in %
			natives (in percent of total)			
no vocational	22	14	9	8	14	-63
vocational	71	75	72	71	73	-2
high school	2	4	7	9	5	440
university	5	7	11	12	8	157
total (in persons)	296,854	325,412	305,229	293,482	310,413	-1
			foreigners (in percent of total)			
no vocational	62	48	37	33	45	-26
vocational	34	46	54	56	48	134
high school	1	2	4	5	3	594
university	3	4	5	6	4	164
total (in persons)	33,675	36,436	48,312	47,162	39,740	40

Source: Authors' calculations based on IABS.

Table 2: The wage-setting curve: IV-estimation results

	$\ln w_{qjk,t-1}$		$\ln u_{qjt}$					
			short run		long run			
education	coeff.	se	coeff.	se	coeff.	se	R^2	obs.
			all experience groups					
all[a)]	0.522	(0.038)***	-0.076	(0.007)***	-0.158	(0.017)***	0.99	768
no vocational	0.587	(0.059)***	-0.063	(0.008)***	-0.152	(0.026)***	0.99	192
vocational	0.593	(0.060)***	-0.077	(0.009)***	-0.188	(0.028)***	0.99	192
high school	0.492	(0.072)***	-0.083	(0.021)***	-0.164	(0.046)***	0.99	192
university	0.445	(0.080)***	-0.091	(0.015)***	-0.164	(0.027)***	0.99	192
			experience group1 (0 – 5 years) and experience group 2 (6 – 10 years)					
no vocational	0.609	(0.083)***	-0.043	(0.009)***	-0.110	(0.040)***	0.99	48
vocational	0.824	(0.058)***	-0.060	(0.008)***	-0.341	(0.102)***	0.99	48
high school	0.819	(0.088)***	-0.050	0.0135***	-0.275	(0.136)**	0.99	48
university	0.434	(0.166)***	-0.123	(0.020)***	-0.218	(0.054)***	0.98	48
			experience group 3 (11 - 15 year) and experience group 4 (16 – 20 years)					
no vocational	0.201	(0.100)**	-0.111	(0.020)***	-0.139	(0.020)***	0.97	48
vocational	0.235	(0.105)**	-0.092	(0.015)***	-0.120	(0.017)***	0.98	48
high school	-0.281	(0.440)	-0.206	(0.092)**	-0.161	(0.043)***	0.79	48
university	0.585	(0.138)***	-0.115	(0.036)***	-0.278	(0.107)***	0.87	48
			experience group 5 (21 - 25 years) and experience group 6 (26 – 30 years)					
no vocational	0.435	(0.115)***	-0.066	(0.013)***	-0.116	(0.025)***	0.97	48
vocational	0.442	(0.217)**	-0.091	(0.027)***	-0.164	(0.070)**	0.86	48
high school	0.658	(0.142)***	-0.081	(0.033)**	-0.237	(0.117)**	0.77	48
university	0.537	(0.130)***	-0.040	(0.019)**	-0.087	(0.036)**	0.85	48
			experience group7 (31 - 35 years) and experience group 8 (> 35 years)					
no vocational	0.364	(0.108)***	-0.095	(0.022)***	-0.149	(0.028)**	0.98	48
vocational	0.228	(0.115)**	-0.067	(0.016)***	-0.087	(0.021)***	0.98	48
high school	0.413	(0.176)**	-0.087	(0.048)*	-0.148	(0.109)	0.65	48
university	0.050	(0.194)	-0.067	(0.028)**	-0.070	(0.023)***	0.91	48

Notes: Dependent variable is $\ln w_{qjt}$, i.e. the log wage in each education-experience group. Errors are heteroskedasticity robust and clustered by education-experience. ***, **, * denote the 1%-, 5%-, and 10%-significance levels. The model is estimated by 2SLS with group specific fixed effects. [a)] A test of overidentifying restrictions cannot reject the null hypothesis that the instruments are valid (Hansen *J*-statistic, *p*-value = 0,61). The Kleibergen-Paap LM statistics rejects the null of underidentification (*p*-value = 0.00) and the Kleibergen-Paap rk Wald F statistic ($F = 83^{***}$) the null that instruments are weak.

Table 3: Partial substitutions elasticities between natives and foreigners, σ , across education-experience cells, ρ, and across education cells, δ

parameter	dependent variable	coefficient	se		observations
σ_{all}	$\ln L_{qjht}/L_{qjft}$	7.01	(0.883)	***	704
σ_{ed1}	$\ln L_{1jht}/L_{1jft}$	3.31	(0.874)	***	704
σ_{ed2}	$\ln L_{2jht}/L_{2jft}$	17.88	(2.414)	***	704
σ_{ed3}	$\ln L_{3jht}/L_{3jft}$	12.96	(5.303)	**	704
σ_{ed4}	$\ln L_{4jht}/L_{4jft}$	2.89	(0.873)	***	704
ρ	$\ln L_{qjt}$	8.57	(1.448)	***	704
δ	$\ln L_{qt}$	2.86	(0.633)	***	84

Notes: Errors are heteroskedasticity robust and clustered by education-experience ***, **, * denote the 1%-, 5%-, and 10%-significance levels. The equations are estimated by 2SLS. Observations are weighted by ln L_{qjt} and ln L_{qt}. The F-test rejects the null hypothesis that all coefficients σ_q are identical across educational groups (F = 29, p-value = 0.00). A test of overidentifying restrictions cannot reject the null hypothesis that the instruments are valid (Hansen J statistic, σ_{all}: p-value = 0.13, ρ: p-value = 0.18; δ: p-value = 0.36). The Kleibergen-Paap LM statistics rejects the null of underidentification (σ_{all}: p-value = 0.00, ρ: p-value = 0.00, δ: p-value = 0.00) and the Kleibergen-Paap rk Wald F statistic (σ_{all}: $F = 56^{***}$, ρ: $F = 21^{***}$, δ: $F = 9^{*}$) the null that instruments are weak.

Table 4: Simulated wage and employment effects of immigration

	1% increase of labor force through immigration						1980-2004 increase (annual average)	
	imperfect labor market				perfect labor market		imperfect labor market	
	short-term effect		long-term effect		short-term	long-term	long-term effect	
	wage	u-rate	wage	u-rate	wage	wage	wage	u-rate
	wages: change in %, unemployment rate: change in %-points							
	total labor force							
all	-0.18	0.31	0.00	0.08	-0.27	0.00	0.00	0.00
no vocational	-0.50	1.01	-0.43	0.52	-0.89	-0.62	0.12	-0.12
vocational	-0.13	0.16	0.06	-0.02	-0.19	0.09	-0.01	0.02
high school	-0.12	0.14	0.08	-0.01	-0.16	0.11	-0.08	0.05
university	-0.11	0.15	0.09	0.01	-0.16	0.11	-0.02	0.03
	native labor force							
all	-0.08	0.09	0.11	-0.06	-0.13	0.15	0.01	0.00
no vocational	-0.03	0.13	0.12	-0.11	-0.12	0.15	0.06	-0.03
vocational	-0.10	0.10	0.09	-0.05	-0.15	0.12	0.00	0.00
high school	-0.08	0.07	0.12	-0.04	-0.12	0.15	-0.05	0.02
university	0.01	0.00	0.23	-0.08	0.01	0.29	0.05	-0.02
	foreign labor force							
all	-1.11	1.97	-1.09	1.16	-1.70	-1.42	-0.09	0.01
no vocational	-1.62	3.06	-1.76	2.00	-2.74	-2.47	0.25	-0.32
vocational	-0.54	0.94	-0.38	0.39	-0.65	-0.38	-0.18	0.22
high school	-0.67	1.08	-0.52	0.40	-0.81	-0.53	-0.61	0.48
university	-2.05	2.50	-2.21	1.44	-3.03	-2.76	-1.16	0.80

Notes: The imperfect labor market scenario is based on our estimates of the wage-setting curves, the perfect labor market scenario assumes that the elasticity of the wage-setting curve approaches infinity and unemployment is zero. The short-term simulations are based on the short-run elasticities of the wage-setting curve and assume that the capital stock remains fixed. The long-run results are based on the long-run elasticities of the wage-setting curve and assume a constant capital-output ratio. In the 1%-scenario, the education and experience composition of the labor supply shock has been taken from the average distribution of the foreign labor force across the education-experience cells during the sample period. The 1980-2004 simulation is based on actual changes of the foreign labor force in each education-experience cell in each year. The annual average is calculated by weighting in each education-experience cell annual changes of wages and of the unemployment rate, respectively, with its share of in the total wage bill and in the labor force,respectively, in each year. Aggregate wage figures are calculated by weighting the wage change of each group by its share in the total wage bill. Aggregate unemployment figures are obtained by weighting each cell with its share in the labor force.

Appendix A1 The explicit solution for the employment response

The general solution for the marginal employment response to an increase in labor supply through immigration is given in equation (6). The model in section 2.2 distinguishes $4 \times 8 \times 2 = 64$ types of labor. Using the notation from the nested production function we write the 1×64 vectors as $\mathbf{x} = [x_{111},\ x_{112}, x_{121}, \dots, x_{211}, \dots, x_{ijk}, \dots, x_{482}]$, where $\mathbf{x} \in \{\mathbf{L}, \mathbf{N}, \mathbf{Y_L}, \mathbf{u}, \boldsymbol{\phi}\}$. The subscript 111 therefore indexes the first, 112 the second, 121 the third, and 482 the 64*th* element of each vector.

Thus, we can write the partial derivative of wages with respect to employment as

$$\frac{\partial \nu^{-1}\mathbf{Y_L}}{\partial \mathbf{L}} = \nu^{-1}\begin{bmatrix} \frac{\partial Y_{L111}}{\partial L_{111}} & \cdots & \frac{\partial Y_{L111}}{\partial L_{ijk}} & \cdots & \frac{\partial Y_{L111}}{\partial L_{482}} \\ \vdots & & \vdots & & \vdots \\ \frac{\partial Y_{Lijk}}{\partial L_{111}} & \cdots & \frac{\partial Y_{Lijk}}{\partial L_{ijk}} & \cdots & \frac{\partial Y_{Lijk}}{\partial L_{482}} \\ \vdots & & \vdots & & \vdots \\ \frac{\partial Y_{L482}}{\partial L_{111}} & \cdots & \frac{\partial Y_{L482}}{\partial L_{ijk}} & \cdots & \frac{\partial Y_{L482}}{\partial L_{482}} \end{bmatrix}. \tag{A.1}$$

Due to the nested structure of the production function we have four types of partial derivatives in equation (A1):

$$\frac{\partial \nu^{-1}Y_{Lqjk}}{\partial L_{qjk}} = \frac{w_{qjk}}{L_{qjk}}\left[s_{qjk}\left\{\frac{1}{\delta} + I^* - \frac{1}{s_q}\left(\frac{1}{\delta} - \frac{1}{\rho}\right) - \frac{1}{s_{qj}}\left(\frac{1}{\rho} - \frac{1}{\sigma_q}\right)\right\} - \frac{1}{\sigma_q}\right],$$

$$\frac{\partial \nu^{-1}Y_{Lqjk}}{\partial L_{ijk'}} = \frac{w_{qjk}}{L_{ijk}}\left[s_{qjk'}\left\{\frac{1}{\delta} + I^* - \frac{1}{s_q}\left(\frac{1}{\delta} - \frac{1}{\rho}\right) - \frac{1}{s_{qj}}\left(\frac{1}{\rho} - \frac{1}{\sigma_q}\right)\right\}\right],$$

$$\frac{\partial \nu^{-1}Y_{Lqjk}}{\partial L_{ij'm}} = \frac{w_{qjk}}{L_{qj'm}}\left[s_{qj'm}\left\{\frac{1}{\delta} + I^* - \frac{1}{s_q}\left(\frac{1}{\delta} - \frac{1}{\rho}\right)\right\}\right],$$

$$\frac{\partial \nu^{-1}Y_{Lqjk}}{\partial L_{q'nm}} = \frac{w_{qjk}}{L_{q'nm}}\left[s_{q'xm}\left\{\frac{1}{\delta} + I^*\right\}\right],$$

where $k \neq k'$, $j \neq j'$, and $q \neq q'$, and s_{qjk}, s_{qj} and s_q denote the share of wages paid to workers in the respective cell of the labor market in the total wage bill. The index function I^* is

$$I^* = \begin{cases} \alpha - 1 & \text{in the short run,} \\ 0 & \text{in the long run,} \end{cases}$$

which follows from the production function if physical capital is fixed in the short run, i.e., if $K = \bar{K}$, and if the capital-output ratio is constant in the long-run, i.e, if $\kappa = \bar{\kappa}$.

Using the wage-setting equation in (3) we can write

$$\frac{\partial \boldsymbol{\phi}}{\partial \mathbf{u}}\frac{\partial \mathbf{u}}{\partial \mathbf{L}} = \begin{bmatrix} \frac{\partial \phi_{111}}{\partial u_{111}}\frac{\partial u_{111}}{\partial L_{111}} & \cdots & 0 & \cdots & 0 \\ \vdots & \ddots & & & \vdots \\ 0 & & \frac{\partial \phi_{ijk}}{\partial u_{ijk}}\frac{\partial u_{ijk}}{\partial L_{ijk}} & & 0 \\ \vdots & & & \ddots & \vdots \\ 0 & \cdots & 0 & \cdots & \frac{\partial \phi_{482}}{\partial u_{482}}\frac{\partial u_{482}}{\partial L_{482}} \end{bmatrix}, \tag{A.2}$$

and

$$\frac{\partial \boldsymbol{\phi}}{\partial \mathbf{u}}\frac{\partial \mathbf{u}}{\partial \mathbf{N}}\frac{d\mathbf{N}}{dM} = \begin{bmatrix} \frac{\partial \phi_{111}}{\partial u_{111}}\frac{\partial u_{111}}{\partial N_{111}}\frac{dN_{111}}{dM} \\ \vdots \\ \frac{\partial \phi_{ijk}}{\partial u_{ijk}}\frac{\partial u_{ijk}}{\partial N_{ijk}}\frac{dN_{ijk}}{dM} \\ \vdots \\ \frac{\partial \phi_{482}}{\partial u_{482}}\frac{\partial u_{482}}{\partial N_{482}}\frac{dN_{482}}{dM} \end{bmatrix}. \tag{A.3}$$

Substituting the matrices (A.1), (A.2), and (A.3) into equation (6) yields the marginal employment response to immigration for the two cases of a fixed capital stock or a constant capital output ratio.

Appendix A2 Sample selection (Western Germany, 1980 - 2004)

	all		natives		foreigners	
	observations	percent	observations	percent	observations	percent
all spells	11,769,882					
- missing nationality	1,661					
spells with known nationality	11,768,221	100.0	10,487,750	100.0	1,280,471	100
- part time workers / trainees	2,543,490	21.6	2,337,020	22.3	206,470	16.1
- age (below 15 and above 60)	166,070	1.4	152,386	1.5	13,684	1.1
- missing education	183,075	1.6	128,945	1.2	54,130	4.2
- wages below social security contribution ceiling	121,748	1.0	109,071	1.0	12,677	1.0
total	8,753,838	74.4	7,760,328	74.0	993,510	77.6

Notes: The IABS includes only wage and salary workers but no self-employed. Due to changes in methodology, allocation of Berlin to the newly formed German states since 1999, and the German unification aggregate statistics for the entire observation period and the region used is not available. During 2000-2004 on average 90 percent of the entire German workforce were dependent employees and 81 percent worked in Western Germany. Thus, about 73 percent of the entire German workforce is covered by our analysis.
Source: Authors' calculations based on the IABS, Statistisches Bundesamt, GENESIS-Online.

Appendix A3 List of variables

GDP: Real GDP (West Germany) at constant 2000 prices. Source: Sachverständigenrat (2009).

Industrial Production: Industrial production incl. construction, volume index (base year: 2000), West Germany. Source: Datastream (code BDIPTOT.G).[16]

CPI: Consumer price index, (base year: 2000), Source: Datastream (code BDCONPRCF).

Export performance: Share of exports in GDP. Source: OECD (2009).

Oil price: Domestic crude oil price, index (base year: 2000). Source: Sachverständigenrat (2009).

Industry mix: The variable is constructed as $indumix_{qj} = \sum_{h=1}^{n} g_{ht} L_{qj\,h,t-1} / L_{qj\,,t-1} - g_t$, where g_{ht} is the employment growth rate in industry h in year t, $L_{qj\,h,t-1}$ is the employment of education-experience group qj in industry h in year $t-1$, $L_{qj\,,t-1}$ is the aggregate employment of education-experience group qj in year $t-1$, and g_t the average overall employment growth rate in year *t*. The summation is over all two digit, non-agricultural, private-sector industries Source: authors' calculations based on IABS.

Export demand: Log of GDP per capita at constant prices and exchange rates of all OECD countries weighted by their average share in German exports during the sample period. Source: authors' calculations based on OECD (2009).

Potential immigration: Estimated immigration potential in each education-experience group. Explanatory variables are GDP per capita at PPP, the unemployment rate and bilateral fixed effects of 20 sending countries. Source: authors' estimates based on immigration stock data from IABS and explanatory variables from the World Bank (2009) and OECD (2009).

Average unemployment benefit: Average unemployment benefit of an unemployed of educational level *q*, experience group *j*, and national origin *k*. Source: IABS.

Ideology index: Share of left- and right-wing parties in government weighted by their seats in parliament. Source: Bjørnskov (2008).

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[16] Access to datasets provided by Datastream was granted through Aarhus University.