Immigration and the Gender Wage Gap*

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Abstract

This paper investigates the effects of immigration on the gender wage gap. Using a detailed individual French dataset, we shed lights on the strong feminization of the immigrant workforce which coincides with an rise in the gender wage gap from 1990 to 2010. Our theoretical model predicts that a shift in the supply of female workers increases the gender wage gap when men and women are imperfect substitutes in production. Our structural estimate points to an imperfect substitutability between men and women workers of similar education, experience and occupation. Our econometric result indicates that a 10% increase in the relative supply of immigrant female workers lowers by 4% the relative wage of female native workers belonging to the same education-experience group. Accounting for cross-group effects, our simulations show that the rise in the relative number of female immigrants decreases the relative wage of female native workers, thereby contributing to a widening native gender wage gap.

Keywords: immigration, wages, gender gap, elasticity of substitution
JEL Classification: F22, J16, J21, J31, J61

*We wish to thank Michel Beine, Isabelle Bensidoun, Frédéric Docquier, Gabriel Felbermayr, Lionel Fontagné, Sébastien Jean, Mariapia Mendola, Gianmarco Ottaviano, Hillel Rapoport as well as participants at the 2016 Royal Economic Society and the 6th International Conference on Economics and Global Interactions for helpful comments. We are grateful to the Centre Maurice Halbwachs (CMH) for having granted us access to the data set. Financial support was provided by the iCODE Institute (Idex Paris-Saclay). Any errors which remain are our own.

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1 Introduction

An extensive literature in social sciences and economics has searched for the driving force behind the gender wage gap. This gap is often related to differences in productivity and discrimination between men and women (Altonji and Blank, 1999; Blau and Kahn, 2000; Goldin, 2014) and also to the role of the feminization of the labor workforce (Topel, 1997; Juhn and Kim, 1999; Acemoglu, Autor, and Lyle, 2004).[1] Surprisingly, the role of immigration has received less attention. It is however a potential important factor as the feminization of the immigration labor force has been one of the most significant trends in the recent years according to the United Nations Population Division (2013).[2] In this paper, we analyze the impact of immigration on the wage gaps between female and male native workers of similar education and experience.

The increasing proportion of employment-related migrants is particularly important in France, where the share of women in the immigrant labor force has increased from 34% in 1990 to 47% in 2010.[3] The feminization of the immigrant workforce might have an impact on the wages of female and male native workers and therefore on the gender wage gap.

We explore this question by using a rich dataset taken from the French labor survey that covers the period from 1990 to 2010. The dataset provides enough detailed information to investigate the impact of immigration on native wages along different important dimensions. In particular, it contains precise information on the educational attainment of natives and immigrants as well as information on their age, occupation and wages.

We show that the evolution of the native wage gap coincides with the increase in the relative number of female immigrants since 1990. We also document significant differences in the pattern of occupation across male and female immigrant workers who share similar education – female immigrants are concentrated in few occupations such as administrative jobs. We also find that occupational differences by gender is more important than occupational differences by nativity status.[4] These last facts may suggest an imperfect substitutability between men and women

[1] In a very interesting study, Acemoglu, Autor, and Lyle (2004) show, for instance, that the supply shift induced by the rapid rise in female labor force participation after World War II has mainly contributed to reduce female wages, thus increasing the gender wage gap.

[2] The United Nations Population Division (2013) reports that about half of the total number of migrants around the world are women. There are few articles investigating the feminization of the migration population, see e.g. the interesting contributions of Zlotnik (1995); Marcelli and Cornelius (2001); Omelaniuk (2005); Docquier, Lowell, and Marfouk (2009); Docquier, Marfouk, Salomone, and Sekkat (2012).

[3] More specifically, we find an increase in the share of female immigrants with higher education. In an recent study, Docquier, Lowell, and Marfouk (2009) examine the skill composition of female migrants. Using the same period of investigation, they find that women represent an increasing share of the OECD immigration stock and exhibit relatively higher skilled emigration rates than men.

[4] Similar results are found in the literature (Anker et al., 1998; Blau, Ferber, and Winkler, 2002; Dustmann, Frattini, and Preston, 2007).

within each education group.

We develop a model that takes into account the possibility of a substitution between male and female workers. In case of an imperfect substitutability between men and women, immigration should impact wages of male and female differently (Topel, 1997). Given *(i)* the strong feminization of the immigrant workforce and *(ii)* the imperfect substitutability between men and women, immigration should increase the relative wage of male native workers. Our results indicate an imperfect substitution between men and women workers with similar education, experience and broad occupational categories.[5] This result is consistent with Acemoglu, Autor, and Lyle (2004) for the U.S. and Pellizzari, Paccagnella, and De Giorgi (2014) for Italy who find that men and women with similar state of residence are imperfect substitutes in production.

We find that an immigration-induced increase in the relative supply of female workers has a negative impact on the relative wage of female native workers. This finding is robust to two complementary empirical methodologies. First, we analyze the effect of the rise in the relative supply of immigrant women on the relative wage of native women workers within the same education-experience group. This econometric analysis which examines the "*within-group*" effects of immigration shows that an immigration-induced increase in the relative labor supply of women raises the wage gap between native men and women with similar education and experience. As the relative labor supply is likely to be endogenous, we implement a set of instrumental variable (IV) regressions. The identification of the causal effect is challenging and rest on the exclusion restriction assumption. As pointed out by Conley, Hansen, and Rossi (2012), it may be more credible to assume that the instrument does not fully satisfy the exclusion restriction. Using their methodology, our main finding remains and is robust to deviations from the exclusion restriction imposed by the IV methodology. Second, we use a structural approach to analyze the overall impact of immigration on the wages of female and male native workers. This methodology allows us to account for the impact of immigrants that are not competing in the same group than native workers (Borjas,

[5]There are many factors which could contribute to the explanation of imperfect substitutability in production between male and female workers of similar education and experience. The gender wage gap might be driven by differences in workers' productive attributes. Borghans, Ter Weel, and Weinberg (2014) show that women are for instance relatively more productive in tasks requiring interpersonal skills, *i.e.*, ability to interact with people (Gilligan, 1982). The prevalence of important gender differences in psychological attributes is an additional source of imperfect substitution between men and women. In particular, women are relatively more risk averse and have lower taste for competition (see the literature review by Croson and Gneezy (2009)). Thus, women tend to sort into occupations with more stable earnings and avoid competitive environments (Marianne, 2011). Gender discrimination in employment should also produce some degree of imperfect substitutability between men and women. Such discriminatory behaviors are supported by Booth and Leigh (2010) who find that females are more likely to find a job when they apply to female-dominated occupations. Another explanation for the existence of the gender gap is the existence of social norms which drive women's decisions to participate in the workforce or induce differential sorting of men and women across occupations (Eccles, 1994; Charles and Grusky, 2005; Marianne, 2011).

2003; Manacorda, Manning, and Wadsworth, 2012; Ottaviano and Peri, 2012). We find a negative effect of the increase in the relative supply of women due to immigration on the wages of women native workers, thereby contributing to a widening gender wage gap.

This article contributes to the immigration literature in several respects. The literature on labor market gender gap documents the differences in jobs and wages between women and men of similar education (Anker et al., 1998; Blau and Kahn, 2000; Goldin, 2014). However, it does not quantify the elasticity of substitution between male and female workers with similar human capital characteristics (*i.e.*, education and experience).[6] We set out a methodology and provide a structural estimate of the elasticity of substitution between male and female workers within the same education and experience group.

While structural methods have been widely applied to samples from different countries, none of the papers use them to investigate the impact of immigration on the relative wage of female native workers.[7] Our model which also incorporates wage rigidities as in D'Amuri, Ottaviano, and Peri (2010) to account for the sluggish adjustment of the French labor market, offers new predictions concerning the effect of immigration on the gender wage gap.

The paper also contributes to a different strand of the literature that focuses on the effect of female migration. While most papers examine the effect of female migration on the labor supply or job specialization of female workers (Amuedo-Dorantes and De La Rica, 2011; Barone and Mocetti, 2011; Cortes and Tessada, 2011; Farré, González, and Ortega, 2011), none have examined the differential effect of immigration on the gender wage gap through its impact on the relative labor supply of women. The literature on the impact of immigration labor supply shocks on wage and employment focuses mostly on male workers. A notable exception is the interesting study of Cortes and Pan (2015) which investigates the relationship between low-skilled immigration and the gender wage gap among high-skilled workers. Using the U.S. intercity variation in low-skilled immigrant flows, they show that a supply shock induced by low-skilled immigrants does not only increase the probability that highly skilled women work long hours, but also leads to a reduction in the gender wage gap in the upper tail of the skill distribution, especially in occupations where the returns to overwork is important. Our econometric analysis examines the effect of immigration on the native wage gap within an education-experience cell. The structural approach takes also

[6] Acemoglu, Autor, and Lyle (2004) and Pellizzari, Paccagnella, and De Giorgi (2014) provide an estimate of the elasticity of substitution between men and women but at the regional level. In this paper, we estimate the degree of substitution between men and women at the education-experience level.

[7] see, e.g., Aydemir and Borjas (2007); Borjas and Katz (2007); Ottaviano and Peri (2012) for the United states, D'Amuri, Ottaviano, and Peri (2010); Felbermayr, Geis, and Kohler (2010); Brücker and Jahn (2011) for Germany, Gerfin and Kaiser (2010) for Switzerland, Manacorda, Manning, and Wadsworth (2012) for the United Kingdom, Edo and Toubal (2015) for France, as well as Docquier, Ozden, and Peri (2013) for OECD countries or Elsner (2013) on the emigration wage effect in Great Britain.

into account the impact on native workers that are not competing in the same skill-cell. In both methodologies, we show that the gender composition of immigrants matters in determining their impact on the gender wage gap.

The remainder of this paper proceeds as follows. In section 2, we describe the theoretical framework used to simulate the overall effects of immigration on wages. In section 3, we describe the data and provide some descriptive statistics. The section 4 provides our estimates regarding the elasticity of substitution between men and women and discusses our choices for the other substitution elasticity values. In section 5, we first provide a regression analysis on the "*within-group*" impact of women and men immigrants on the relative wage of native women at the education-experience level. We then impose our theoretical structure to simulate our model and quantify the impact of immigration on the wage of female and male native workers. The last section concludes.

2 Theoretical Framework

This paper uses two complementary strategies to investigate the impact of immigration on the wages of native men and women. First, we use a "non-structural approach" to estimate how the gender wage gap within education-experience groups is affected by the feminization of the immigrant workforce experienced by that group. Second, we investigate the full wage response of native men and women to immigration by accounting for the cross-effects of immigration on the wages of complementary workers and capital accumulation. We therefore need to impose a structure on the technology of the labor market and follow the well-established literature on the wage response to labor supply and demand shocks (Katz and Murphy, 1992; Card and Lemieux, 2001; Borjas, 2003). We extend this literature by taking into account the substitutability between men and women workers.

The nested Constant Elasticity of Substitution structure. In theory, immigration should reduce the wages of competing workers (who have skills similar to those of the migrants), and increase the outcomes of complementary workers (who have skills that complement those of immigrants). In the medium- and long-run, firms should moreover respond to the increased supply of immigrants through capital accumulation to compensate the fall of the capital-labor ratio. In order to account for the various effects of immigration on native wages, we combine different skill cells in a multi-stage nested Constant Elasticity of Substitution (CES) production function (Borjas, 2003; Manacorda, Manning, and Wadsworth, 2012; Ottaviano and Peri, 2012).

We assume a constant-return-to-scale aggregate production function that takes the Cobb-Douglas form (D'Amuri, Ottaviano, and Peri, 2010; Brücker and Jahn, 2011; Ottaviano and Peri, 2012; D'Amuri and Peri, 2014).[8] The aggregate production function is given by Equation (1). The physical capital K_t and a labor composite L_t are combined to produce output Y_t at time t.

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

where A_t is exogenous total factor productivity (TFP) and $\alpha \in (0,1)$ is the income share of labor. L_t is defined as a composite of different categories of workers who have different level of education, work experience and nativity. We follow the literature on the wage structure (Katz and Murphy, 1992; Autor, Katz, and Kearney, 2008; Goldin and Katz, 2009) or on migration (Card and Lemieux, 2001; Card, 2009; Ottaviano and Peri, 2012; Docquier, Ozden, and Peri, 2013) by assuming L_t to have a nested CES structure, which combines the labor supply of two broad education groups $b \in \{H, L\}$. L_{Ht} and L_{Lt} are aggregate measures of the labor supply of high and low educated workers, respectively.

$$L_t = \left[\theta_{Ht} \cdot L_{Ht}^{\rho_{HL}} + \theta_{Lt} \cdot L_{Lt}^{\rho_{HL}}\right]^{1/\rho_{HL}} \tag{2}$$

The parameters θ_{Ht} and θ_{Lt} measure the relative efficiency of each category, with $\theta_{Ht} + \theta_{Lt} = 1$. $\rho_{HL} = (\sigma_{HL} - 1)/\sigma_{HL}$ with σ_{HL} being the degree of substitution between the group of high educated workers and the group of low educated workers.

As in Docquier, Ozden, and Peri (2013), the education groups L_{Ht} and L_{Lt} respectively refer to an education level beyond high school and less than college. Implicitly, this classification assumes that within these two large groups, workers with different levels of education are perfect substitutes. As noted in Manacorda, Manning, and Wadsworth (2012), a classification that uses only two education groups might thus be too restrictive. More generally, the number of education groups that has to be used is subject to controversy (Card, 2012; Borjas, Grogger, and Hanson, 2012). In order to give more flexibility to our model and test whether a two education grouping is relevant in the French context, we therefore split up L_{Ht} and L_{Lt} into two finer education groups $j \in \{1, 2\}$ as follows:

$$L_{Ht} = \left[\theta_{H_1 t} \cdot L_{H_1 t}^{\rho_H} + \theta_{H_2 t} \cdot L_{H_2 t}^{\rho_H}\right]^{1/\rho_H} \tag{3}$$

[8] This assumption allows us to concentrate exclusively on the long-run effects of immigration on wages, by considering full capital adjustment as a response to immigrant-induced labor supply shifts.

$$L_{Lt} \quad = \quad \left[\theta_{L_1t} \cdot L_{L_1t}^{\rho_L} + \theta_{L_2t} \cdot L_{L_2t}^{\rho_L}\right]^{1/\rho_L} \tag{4}$$

where $\rho_H = (\sigma_H - 1)/\sigma_H$ and $\rho_L = (\sigma_L - 1)/\sigma_L$, with σ_H and σ_L capture the degree of substitution between workers within each broad education groups. The terms L_{bjt} for $bj \in \{H_1, H_2, L_1, L_2\}$ are aggregate measures of labor supplied by workers with, respectively, some college education (H_1), a college degree (H_2), less than high school education (L_1) and a high school education (L_2). For ease of clarification, we note $\sigma_b \in \{\sigma_H, \sigma_L\}$. The parameters θ are the education specific productivity levels with $\theta_{H_1t} + \theta_{H_2t} = 1$ and $\theta_{L_1t} + \theta_{L_2t} = 1$.

As shown by Edo and Toubal (2015), the education group L_{Ht} in France represents 25% of the total number of workers and is composed of homogeneous individuals – *i.e.*, the workers with some college (L_{H_1t}) or a college degree (L_{H_2t}) are perfect substitutes in production. As a result, we consider that $\sigma_H \to \infty$.[9] However, within the group of low educated workers,[10] Edo and Toubal (2015) find that workers with high school education (L_2) and less than a high school education (L_1) are imperfect substitutes (with an elasticity of substitution equals to 10). We thus keep the decomposition of the low educated group L_{Lt} into two finer education classes, denoted L_{L_1t} and L_{L_2t}.[11] As a result, we use two broad education groups $b \in \{H, L\}$, and three education classes grouping together workers with tertiary education L_{Ht}, secondary education L_{L_2t} and primary education – where $L_{bjt} \in \{L_{L_1t}, L_{L_2t}, L_{Ht}\}$.[12]

Considering the experience level of workers, we divide the labor composite L_{bjt} into four experience intervals of five years [1-10 ; 10-20 ; 20-30 ; 30-40], as in Felbermayr, Geis, and Kohler (2010); Gerfin and Kaiser (2010); Elsner (2013); Brücker, Hauptmann, Jahn, and Upward (2014). This strategy has three main advantages. First, it attenuates the impact of any potential bias regarding our experience measure, and in particular, the fact that employers may evaluate the experience of immigrants differently than natives. Second, by increasing the number of observation per cell, our strategy with 12 skill-cells (three education groups and four experience groups) tends to correct for the attenuation bias our estimates may suffer (Aydemir and Borjas, 2011). Third,

[9]This follows Card and Lemieux (2001); Card (2009). For the United States, they do not disaggregate the high education group corresponding to college equivalent workers.

[10]The low educated category L_{Lt} regroups 75% of workers, among which 62.6% have a high school degree and 37.4% have an education below high school (Edo and Toubal, 2015).

[11]This disaggregation is consistent with Goldin and Katz (2009); Borjas, Grogger, and Hanson (2012); Ottaviano and Peri (2012).

[12]D'Amuri, Ottaviano, and Peri (2010); Gerfin and Kaiser (2010); Elsner (2013) also use three education classes to investigate the wage impact of migration in Germany, Switzerland and Lithuania, respectively. In addition, our classification using three education groups is consistent with the study by Brücker, Hauptmann, Jahn, and Upward, (2014, p. 211) which shows that "it is most suitable to distinguish three education groups in European labor markets."

since women tend to face more frequent periods of inactivity or unemployment, the correspondence between their potential and effective experience may collapse. It is therefore relevant to use four but broader experience groups to study the labor market effects of immigration on male and female natives.

$$L_{bjt} = \left[\sum_{k=1}^{4} \theta_{bjkt} \cdot L_{bjkt}^{\rho_X}\right]^{1/\rho_X} \tag{5}$$

L_{bjkt} is the number of workers with education bj and experience k at time t. We define $\rho_X = (\sigma_X - 1)/\sigma_X$, where σ_X measures the elasticity of substitution across the different experience classes and within a narrow education group. The parameters θ_{bjkt} capture the relative efficiency of workers within the education-experience group.

We allow men and women be imperfect substitutes. As they work in different jobs (Anker et al., 1998; Blau, Ferber, and Winkler, 2002), they may not compete in the same labor market. One reason behind this labor market segmentation may lie in important gender differences in attitudes, behaviors and other types of productive characteristics (see footnote 5). We thus assume that L_{bjkt} incorporates contributions of workers who differ in gender, and we divide L_{bjkt} into a CES aggregator of males and females as follows:

$$L_{bjkt} = \left[\theta_{S_M bjkt} \cdot Male_{bjkt}^{\rho_F} + \theta_{S_F bjkt} \cdot Female_{bjkt}^{\rho_F}\right]^{1/\rho_F} \tag{6}$$

where $\rho_F = (\sigma_F - 1)/\sigma_F$ and σ_F denotes the elasticity of substitution between men and women, while $\theta_{S_M bjkt}$ and $\theta_{S_F bjkt}$ stand for their specific efficiency levels standardized so that $\theta_{S_M bjkt} + \theta_{S_F bjkt} = 1$. We estimate the parameter σ_F in section 4.

Turning to the nativity of workers, we follow Manacorda, Manning, and Wadsworth (2012); Ottaviano and Peri (2012) and assume L_{Sbjkt} to be a CES aggregate of native-born N_{Sbjkt} and of foreign-born workers M_{Sbjkt}:

$$L_{Sbjkt} = \left[\theta_{Sbjkt}^{N} \cdot N_{Sbjkt}^{\rho_I} + \theta_{Sbjkt}^{M} \cdot M_{Sbjkt}^{\rho_I}\right]^{1/\rho_I} \tag{7}$$

where $\rho_I = (\sigma_I - 1)/\sigma_I$ and σ_I captures the degree of substitution between natives and immigrants in an education-experience cell. The relative efficiency for each group of workers is given by the productivity parameters θ_{Sbjkt}^{N} and θ_{Sbjkt}^{M}, with $\theta_{Sbjkt}^{N} + \theta_{Sbjkt}^{M} = 1$. We estimate the parameter σ_I in section 4.

The assignment of natives and immigrants to skill-cells may be inaccurate if immigrants are more likely to accept jobs requiring lower education than they have (Dustmann, Frattini, and Preston, 2013). We may overestimate the immigrant contribution to the supply of high educated workers and underestimate the immigrant contribution to the supply of low educated workers. As a result, the estimates of σ_I and the estimated wage effects of immigration may not be fully correct. In France, however, Docquier, Ozden, and Peri (2013) find very weak evidence of educational downgrading among immigrants: highly educated immigrants are as likely to be in highly skilled occupation as natives. This result is consistent with a report by the OECD (2007) showing that France is one of the OECD country with the lowest rate of downgrading among immigrants (see Tables II.2 and II.A3.1 of the report). Hence, accounting for the potential downgrading of immigrants in France does not alter the simulation results of Docquier, Ozden, and Peri (2013) who quantify the wage effect of immigration.

Introducing wage rigidities. The nested CES framework allows to compute the overall impact of immigration under the assumption that wages adjust perfectly to labor supply shocks. However, the French labor market characterized by high minimum wage, high unemployment benefits, strict employment protection, powerful labor unions and product market rigidities (e.g., firms entry cost) may prevent full wage adjustment (Angrist and Kugler, 2003). These dimensions should affect the wage-setting mechanism (Babecký, Du Caju, Kosma, Lawless, Messina, and Rõõm, 2010), the reservation wage (Cohen, Lefranc, and Saint-Paul, 1997) and the scope for bargaining, which in turn should have an impact on the responsiveness of wages to labor supply shocks (as suggested by D'Amuri, Ottaviano, and Peri (2010); Felbermayr, Geis, and Kohler (2010); Brücker and Jahn (2011); Brücker, Hauptmann, Jahn, and Upward (2014)).[13] As a result, wages should not adjust perfectly to immigrant-induced supply shifts in France.

We follow the methodology of D'Amuri, Ottaviano, and Peri (2010) and extend the structural approach by accounting for wage rigidities. In their model, they assume that "a change in wages [produced by immigration] may induce an employment response for natives" (D'Amuri, Ottaviano, and Peri (2010), p. 553). The total effect of the supply shift induced by immigration on native wages can thus be decomposed into *(i)* a direct wage effect due to immigrant-induced supply shifts and *(ii)* an indirect wage effect through the extensive margin of the labor supply. In our model, the indirect wage effects thus attenuate the direct wage effects induced by immigration, allowing for a sluggish adjustment of wages.

We follow Edo (2015) who use similar data for France over the 1990-2002 period and assume

[13] This is also consistent with Card, Kramarz, and Lemieux (1999) who show that labor supply shocks have less impact on the adjustment of wages in France because of wage rigidities.

that a 10% increase in the immigrant population reduce the native workers' population by 3% with similar experience and education.[14] This approach does not require any assumption about the type of rigidities that should impact the responsiveness of wages to immigration. In the simulations, we therefore impose an employment reaction for native workers in response to immigration so as to allow wages to adjust imperfectly. Such pattern is consistent with the prevalence of rigid labor market institutions (Angrist and Kugler, 2003; Edo, 2016).

Labor market equilibrium. In equilibrium, wages (and employment) levels are such that firms maximize profits. Profit-maximizing firms pay each skill group a real wage equal to the group's marginal product. By equating native wage to the marginal products of their labor ($w^N_{Sbjkt} = \delta Y_t/\delta N_{Sbjkt}$) and using Equations (1) to (7), we can derive an expression for the equilibrium wage of natives for each gender-education-age-time cell:

$$\begin{aligned} log\left(w^N_{Sbjkt}\right) \;=\; & log\left(\alpha \cdot A_t \cdot [\kappa_t]^{1-\alpha}\right) + \frac{1}{\sigma_{HL}} \cdot log\left(L_t\right) + log\left(\theta_{bt}\right) - \left[\frac{1}{\sigma_{HL}} - \frac{1}{\sigma_b}\right] log\left(L_{bt}\right) \\ & + log\left(\theta_{bjt}\right) - \left[\frac{1}{\sigma_b} - \frac{1}{\sigma_X}\right] log\left(L_{bjt}\right) + log\left(\theta_{bjkt}\right) - \left[\frac{1}{\sigma_X} - \frac{1}{\sigma_I}\right] log\left(L_{bjkt}\right) \\ & + log\left(\theta_{Sbjkt}\right) - \frac{1}{\sigma_F} \cdot log\left(L_{Sbjkt}\right) + log\left(\theta^N_{Sbjkt}\right) - \frac{1}{\sigma_I} \cdot log\left(N_{Sbjkt}\right) \end{aligned} \quad (8)$$

$\kappa_t = (K_t/L_t)$ and the productivity parameters θ measure the specific productivity levels between workers with different education, experience, gender and nativity. σ_{HL}, σ_b, σ_X, σ_F and σ_I respectively measure the elasticities of substitution between the high and low education groups, within both broad education groups, between experience groups, between men and women and between natives and immigrants.

Any changes in one of the factors on the right-hand side affect the marginal product, which leads to a change in the real wage *ceteris paribus*. According to Equation (8), an immigration-induced supply shift in a given skill group S, bj, and k thus generates a (potentially negative) partial effect on the wages of native workers in the same gender-education-experience group, as well as (potentially positive) cross-effects on the wages of natives in other groups.

Wage effects of immigration. The overall impact of immigration on native wages can be derived from the demand function (8). As in Borjas (2003); Ottaviano and Peri (2012), the TFP and productivity levels are assumed to be insensitive to immigration. Thus, we can express the percentage wage changes due to immigrants for natives in the long-run as:

[14] This magnitude is found by Glitz (2012) for Germany.

$$
\begin{aligned}
\left(\frac{\triangle w^N_{bjkt}}{w^N_{bjkt}}\right) = & \left[\frac{1}{\sigma_{HL}}\right]\sum_b\sum_j\sum_k\sum_S\left(s^M_{Sbjkt}\cdot\frac{\triangle M_{Sbjkt}}{M_{Sbjkt}}+s^N_{Sbjkt}\left(\frac{\triangle N_{Sbjkt}}{N_{Sbjkt}}\right)_{response}\right)\\
& -\left[\frac{1}{\sigma_{HL}}-\frac{1}{\sigma_b}\right]\left(\frac{1}{s_{bt}}\right)\sum_j\sum_k\sum_S\left(s^M_{Sbjkt}\cdot\frac{\triangle M_{Sbjkt}}{M_{Sbjkt}}+s^N_{Sbjkt}\left(\frac{\triangle N_{Sbjkt}}{N_{Sbjkt}}\right)_{response}\right)\\
& -\left[\frac{1}{\sigma_b}-\frac{1}{\sigma_X}\right]\left(\frac{1}{s_{bjt}}\right)\sum_k\sum_S\left(s^M_{Sbjkt}\cdot\frac{\triangle M_{Sbjkt}}{M_{Sbjkt}}+s^N_{Sbjkt}\left(\frac{\triangle N_{Sbjkt}}{N_{Sbjkt}}\right)_{response}\right)\\
& -\left[\frac{1}{\sigma_X}-\frac{1}{\sigma_F}\right]\left(\frac{1}{s_{bjkt}}\right)\sum_S\left(s^M_{Sbjkt}\cdot\frac{\triangle M_{bjkt}}{M_{bjkt}}+s^N_{Sbjkt}\left(\frac{\triangle N_{Sbjkt}}{N_{Sbjkt}}\right)_{response}\right)\\
& -\left[\frac{1}{\sigma_F}-\frac{1}{\sigma_I}\right]\left(\frac{1}{s_{Sbjkt}}\right)\left(s^M_{Sbjkt}\cdot\frac{\triangle M_{Sbjkt}}{M_{Sbjkt}}+s^N_{Sbjkt}\left(\frac{\triangle N_{Sbjkt}}{N_{Sbjkt}}\right)_{response}\right)\\
& -\left[\frac{1}{\sigma_I}\right]\left(\frac{\triangle N_{Sbjkt}}{N_{Sbjkt}}\right)_{response}+(1-\alpha)\left(\frac{\triangle\kappa_t}{\kappa_t}\right) \qquad (9)
\end{aligned}
$$

where s_{bt}, s_{bjt}, s_{bjkt}, s_{Sbjkt}, s^M_{Sbjkt} and s^N_{Sbjkt} are the shares of the total wage income paid to the respective groups.[15] The terms $(\triangle M_{Sbjkt}/M_{Sbjkt})$ and $(\triangle N_{Sbjkt}/N_{Sbjkt})_{response}$ are the changes in immigrant and native labor supply in the same respective groups over the corresponding period.

The fraction $(\triangle M_{Sbjkt}/M_{Sbjkt})$ represents the percentage change in the supply of immigrant workers with gender S, education bj and experience k between 1990 and 2010. The terms capture the direct effect of the change in the supply of immigrants on wages. The terms associated with the subscript "response" account for employment effects caused by the change in the supply of immigrant workers in a given skill-cell – *i.e.,* $(\triangle N_{Sbjkt}/N_{Sbjkt})_{response}$ represents the change in labor supply of native workers in the same group caused by immigration. These terms capture the indirect wage effects due to the change in native employment caused by immigrants. As a result, the percentage wage changes due to immigration depends on *(i)* the income shares accruing to the various factors, *(ii)* both direct and indirect wage effects depending on the size of the immigration supply shock and *(iii)* the various elasticities of substitution that lie at the core of the CES framework.

As in D'Amuri, Ottaviano, and Peri (2010), the response of native employment in the cell (S, b, j, k, t) is obtained by multiplying 0.7 by the change in immigrants standardized by the initial employment – *i.e.* $(\triangle M_{Sbjkt}/M_{Sbjkt} + N_{Sbjkt})$. An inflow of immigrants equal to 1% of the initial

[15] For instance, $s^M_{Sbjkt} = \left(w^M_{Sbjkt}M_{Sbjkt}\right)/\sum_b\sum_j\sum_k\sum_S\left(w^M_{Sbjkt}M_{Sbjkt}+w^N_{Sbjkt}N_{Sbjkt}\right)$ is the share of total wage income in period t paid to migrant workers with with gender S, education bj and experience k.

employment in a skill group increases total employment in that group by 0.7%. This strategy assumes that the displacement effects due to immigration are similar across cells. Section 5 discusses the implications of this assumption.

Using the percentage change in wages for each cell (S, b, j, k, t), we can then aggregate and find the effect of immigration on several representative wages. From Equation (9), we can calculate the mean wage effect of immigration for the various education groups by taking the weighted average of the wage effects across the experience groups for a particular education group. The weights are given by income shares. By taking the weighted average of the wage effects across education groups, we can compute the total wage effects due to immigration.

3 Data and Facts

Sample and variables. The analysis uses data from the French annual labor force survey (LFS) which covers 21 years of individual-level data for the period 1990 to 2010. The LFS is conducted by the French National Institute for Statistics and Economic Studies (INSEE).[16] It provides detailed information on demographic and social characteristics at the individual level. Our data are composed of two distinct waves that we merge together. The first wave of data (1990-2002) covers a random sample of around 145,000 individuals per year, with a sampling rate equal to 0.35%. The second wave of data (2003-2010) covers a random sample of around 295,000 individuals per year, the sampling rate is equal to 0.65%.

In order to make our sample representative of the French population, we systematically use an individual weight (computed by the INSEE). This weight indicates the number of individuals each observation represents in the total population.[17]

Our empirical analysis uses information on individuals aged from 16 to 64. We follow several studies on immigration and we drop all individuals who are self-employed (such as farmers, business owners, liberal professions), in military occupations or enrolled at school (Borjas, 2003, 2006; Ottaviano and Peri, 2012). We also exclude all persons in the clergy since the wage-setting mechanism in that occupation should differ from the wage-setting of all other workers.

We define an immigrant as a person born outside France and who is either a noncitizen or a naturalized citizen. All other individuals are considered as natives. This definition is similar to (Borjas, 2003, 2006; Ottaviano and Peri, 2012). In this paper, we use the terms immigrant and foreign-born interchangeably.

[16] Institut National de la Statistique et des Etudes Economiques.

[17] The weight is called "extri". For the period 2003-2010, we follow the suggestion by the INSEE and we use "extri16" as weight to compute average wage by skill group.

The employment survey divides the education level into six categories: college graduate, some college, high school graduate, some high school, just before high school, no education. According to the International Standard Classification of Education (ISCED), those levels of education respectively correspond to *(1)* a second stage of tertiary education, *(2)* first stage of tertiary education, *(3)* post-secondary non-tertiary education, *(4)* (upper) secondary education, *(5)* lower secondary education and *(6)* a primary or pre-primary education. In order to classify the workers in terms of their level of education, we merge the two highest education groups *(1)* & *(2)* to build the group L_H, the two medium *(3)* & *(4)* and the two lowest *(5)* & *(6)* to build L_{L_2} and L_{L_1}, respectively.

Individuals with the same level of education, but a different age or experience are unlikely to be perfect substitutes (Card and Lemieux, 2001). Hence, individuals are characterized by their labor market experience. Following Mincer (1974), work experience is computed by subtracting for each individual the age of schooling completion from reported age. This measure differs from the one used in the migration literature since the age of completion of schooling is usually unavailable.[18] For a few surveyed individuals, the age of completion of schooling is very low, between 0 and 11 inclusive. Since individuals cannot start accumulating experience when they are too young, we have raised the age of completion of schooling for each surveyed individual to 12 if it is lower. We restrict the analysis to persons who have between 1 and 40 years of experience

The survey reports for each worker the monthly wage net of employee payroll tax contributions adjusted for non response, as well as the usual number of hours worked a week.[19] Since wages are reported in nominal terms, we deflate the data using the French Consumer Price Index provided by the INSEE. Based on monthly wages and hours worked, we compute the mean log hourly wage as a proxy for the price of labor at the skill-cell level.

In order to compute the substitution elasticities and to simulate our model, we need a proxy for labor supply in each skill group. From a theoretical point of view, the labor quantity in a specific cell stands for the number of "efficiency units" provided by all workers. Unless otherwise specified, we follow Borjas, Grogger, and Hanson (2008, 2011); D'Amuri, Ottaviano, and Peri (2010) and express labor supply as the level of total employment in a specific cell. This strategy is also consistent with Manacorda, Manning, and Wadsworth (2012) who use population as a measure of labor supply.

[18]Empirical works rather assign a particular entry age into the labor market to the corresponding educational category.

[19]For the period 1990-2002, we use the number of hours worked during the previous week for those workers who do not have regular and fixed hours worked.

The gender distribution of immigrants over time. In France, the stock of female immigrants has grown proportionally faster than the stock of male immigrants. In this regard, Figure 1 shows that the share of female immigrant workers in France has increased more than the share of male immigrant workers, going from 34% of the immigrant labor force in 1990 to 47% in 2010.[20]

Since 2008, women immigrants represent 51% of the total immigrant population in France (Beauchemin, Borrel, and Régnard, 2013). The share of female immigrants in the total labor force has doubled, increasing from 2.3% in 1990 to 4.3% in 2010, while the share of male immigrants only increased from 4.5% to 4.9%.[21]

– Insert Figure 1 about here –

The educational distribution of natives and immigrants by gender. Table 1 shows the educational composition of natives and immigrants in the labor force by gender groups. For both males and females, we display the fraction of natives and foreign-born in the labor force for the high, medium and low education categories in 1990 and 2010. Table 1 shows that the labor force becomes increasingly educated over time. In particular, the table illustrates the rapid increase of the fraction of immigrants in the highly educated category (with some college or more). Irrespective of gender, about 28% of immigrants are classified as highly educated, against less than 10% in 1990. However, the growth in high educated labor force is faster for immigrants than for native workers. This partly reflects the shift toward a migration policy which is more selective (Edo and Toubal, 2015). Among the population of immigrants, the educational shift is more pronounced for women than for men workers.

– Insert Table 1 about here –

Gender, migration and occupation. In Table 2, we report the average share of workers according to their gender and origin by occupation. We group workers into five broad occupational categories defined according to the socio-professional categorization of INSEE. The table reports first strong differences in occupation by gender (as in e.g., Anker et al. (1998); Blau, Ferber, and Winkler (2002); Jurajda (2003); Collado, Iturbe-Ormaetxe, and Valera (2004)). While female workers constitute more than 50% of administrative workers, male workers represents less than 15% of this socio-professional category. Male workers are concentrated in skilled and unskilled

[20]We do not follow workers across years in the French annual labor force survey. We are not able to explain whether the shift in the immigrant work force comes from existing female immigrants entering the labor market or new immigrants with higher labor force participation rates.

[21]In terms of workforce participation, the share of female immigrants in the female labor force went from 5% to 9%, while the share of male immigrants in the male labor force went from 8% to 10%.

manual occupations. This finding are in line with Dustmann, Frattini, and Preston (2007) that show a disproportionate concentration of women in intermediate non-manual occupations and in personal services in the United Kingdom. Sikora and Pokropek (2011) also show that women are leading men in their interest in non-manual occupations in almost all countries.

– Insert Table 2 about here –

We also find that immigrants tend to be concentrated in skilled and unskilled manual jobs as compared to natives. Using data on immigration in the United Kingdom, Dustmann, Frattini, and Preston (2013) find a stronger concentration of immigrants in unskilled jobs. Importantly, Table 2 indicates that the occupational differences are more important between men and women than between natives and immigrants. Similar results are found for the United Kingdom (Dustmann, Frattini, and Preston, 2007) and Spain (Amuedo-Dorantes and De La Rica, 2011).[22]

In Table 3, we cross the information on education, gender and occupation. We find strong occupational differences between similarly educated men and women which suggests that they may compete for different types of jobs. Men and women might therefore be imperfect substitutes in production.

– Insert Table 3 about here –

The occupational differences between men and women may be an indicator of their imperfect substitutability if these two groups have different productive characteristics. If men and women of similar education have distinct characteristics, they should be in different occupations (as shown in Table 3) competing for different types of jobs. In addition, men and women of similar education and experience may be considered as imperfect substitutes by employers due to gender discrimination. Such behaviors among employers could produce some imperfect substitutability between men and women (even if they are equally productive), leading to gender occupational differences as observed in Table 3.

Conditional gender wage gap and immigrant induced supply shift. We use the individual level data to estimate the gender wage gap and its yearly evolution. We apply standard Mincer regressions and estimate yearly cross-sections of wages' determinants controlling for gender and a wide set of education and experience specific effects. We then contrast the estimated coefficients with the shift in the structure of the immigrant work force. The details and results of the estimation are provided in Appendix.

[22] It is also documented that occupational segregation by gender in the United States is more important than occupational segregation by race or ethnicity (Blau, Ferber, and Winkler, 2002).

– Insert Figure 3 about here –

Figure 3 displays the evolution of the estimated gender wage gap (left axis) and the yearly ratio of female to male immigrants workers (right axis).[23] The figure shows a strong and positive correlation between the estimated wage gap and the increase in the relative number of female immigrants (the correlation coefficients between the two series is about 0.83). This descriptive result suggests that the supply shift due to female immigrants has probably affected the relative wage of female native workers. Such relationship would be consistent with the fact that men and women of similar education and experience tend to be imperfect substitutes.

4 Elasticities of Substitution

In this section, we derive two empirical equations from our structural model to estimate the structural parameters σ_F (*i.e.*, the elasticity of substitution between men and women) and σ_I (*i.e.*, the elasticity of substitution between immigrants and natives). As our model accounts for wage rigidities, the total effect of the supply shift induced by immigration on native wages can be decomposed into *(i)* a direct wage effect due to immigrant-induced supply shifts and *(ii)* an indirect wage effect through the extensive margin of the labor supply. In our model, the indirect wage effects thus attenuate the direct wage effects induced by immigration. This indirect effect could be gender specific as the labor supply of women tend to be more elastic (see, e.g., Blau and Kahn (2007) as well as the meta-analysis by Evers, De Mooij, and Van Vuuren (2008)). This implies an adjustment via the extensive margin that is more pronounced for women than for men. In order to determine the elasticity of substitution, σ_I and σ_F, we are examining the total net effect of the supply shift and do not decompose it into his direct and indirect components.

The elasticities of substitution between education and experience groups are taken from Edo and Toubal (2015) who already estimate them for France over our period of interest (1990-2010).[24]

Elasticity of substitution between women and men. In order to estimate the elasticity of substitution between men and women, we examine how their relative supply changes line up with their relative wage changes. For the analysis, we divide our sample into 24 education and

[23]Notice that the relative number of female to male immigrant workers increased for our three education groups (high, medium, low) over the 1990-2010 period. In fact, this share increased from 47% in 1990 to 105% in 2010 for the high educated group, from 42% in 1990 to 81% in 2010 for the medium educated group, and from 48% in 1990 to 78% in 2010 for the low educated group.

[24]The estimated values of the elasticities of substitution between education and experience groups do not depend on the nested CES structure below these levels of aggregation (*i.e.*, do not depend on σ_F and σ_I).

experience cells for each year of our sample period. For each skill group, we compute the relative hourly wage of women, $\left(w_{bjkt}^{Female}/w_{bjkt}^{Male}\right)$ and their relative labor supply, $(Female_{bjkt}/Male_{bjkt})$.

In Figure 2, we provide a preliminary look at the correlation between the relative supply of women and their relative hourly wages for the sample of full-time workers.[25] Figure 2 indicates a negative and significant relationship between relative hourly wages and relative labor supplies across skill groups. The unconditional correlation between the relative wages and the relative supplies is about 0.05 (with a corresponding t-student equal to -4.24) which suggests an elasticity of substitution σ_F of about 20. This finding provides a first evidence of the imperfect substitutability between men and women.

– Insert Figure 2 about here –

In Figure 2, we implicitly assume a constant relative demand for women. This demand may however differ across skill groups and over time. We thus need to estimate the impact of the relative supply of women on their relative wages controlling for the set of education, experience, time fixed effects, and their interactions. The estimation will therefore take into account the systematic changes of female workers' relative productivity at the education level and experience level. In Equation (10), we consider the log of the relative wage of women in a particular skill group with the specification,

$$log\left(\frac{w_{bjkt}^{Female}}{w_{bjkt}^{Male}}\right) = log\left(\frac{\theta_{S_F bjkt}}{\theta_{S_M bjkt}}\right) - \frac{1}{\sigma_F}\cdot log\left(\frac{Female_{bjkt}}{Male_{bjkt}}\right) \tag{10}$$

where w_{bjkt}^{Female} and w_{bjkt}^{Male} are the wages of women and men workers in cell (bj, k, t), respectively. In order to compute the dependent variable, we follow the standard approach and use the mean of log wages (Katz and Murphy, 1992; Borjas, 2003; Borjas, Grogger, and Hanson, 2012). The explanatory variable is the log relative number of female workers in employment. The term $log\left(\theta_{S_F bjkt}/\theta_{S_M bjkt}\right)$ captures the relative productivity of women. In order to estimate the elasticity of substitution between men and women (denoted σ_F), we assume that the relative productivity term can be captured by a vector of fixed effects δ_{bjkt} and a group-specific error term ξ_{bjkt} uncorrelated with the log relative supply of women. Because the dependent variable is the ratio of female over male log hourly wages, factors that affect symmetrically female and male labor demand are removed from the equation.

In columns 1 and 2 of Table 4, we report the estimated coefficients from Equation 10. We control for the complete set of education, experience and time fixed effects. We follow Borjas,

[25] The inclusion of part-time workers in the sample does not change our results and conclusions.

Grogger, and Hanson (2012, p. 204) and weight the observations by the inverse variance of the ratio of the log wages for each education-experience-time cell. We cluster the standard errors at the education-experience level to allow error correlation within skill group.

The estimated coefficients are negative and significant for the full-time sample and the full- and part-time sample, suggesting an imperfect substitution between men and women – our estimated coefficients imply an substitution elasticity between 12.5 and 14.[26]

– Insert Table 4 about here –

In column 3, we slice each education-experience group into occupational categories. Our dataset provides information on five occupations categories: professionals and managers, supervisors, administrative workers, skilled-manual workers and unskilled-manual workers. We thus estimate the elasticity of substitution between men and women at the *education-experience-occupation* level. We control for a complete set of education, experience, occupation and time effects, and we cluster the standard errors at the occupation-education-experience to account for within-cluster correlation. Our baseline estimate, reported in column 3, indicates a degree of substitutability by around 16.7. Within occupations, men and women of similar education and experience are still imperfect substitutes, although their degree of substitution is higher than in columns 1 and 2. This result is consistent with the literature on gender explaining that men and women do not only sort into different occupations, they also perform different set of tasks within occupations (see footnote 5).

Elasticity of substitution between natives and immigrants. We structurally derive an equation that allows us to estimate the degree of substitutability between immigrants and natives:

$$log\left(\frac{w^M_{Sbjkt}}{w^N_{Sbjkt}}\right) \quad = \quad log\left(\frac{\theta^M_{Sbjkt}}{\theta^N_{Sbjkt}}\right) - \frac{1}{\sigma_I} \cdot log\left(\frac{M_{Sbjkt}}{N_{Sbjkt}}\right) \tag{11}$$

As compared to Equation (10), this equation has four main dimensions: education, experience, time and *gender*. The dependent variable is the relative log mean wage of immigrants in a particular skill-gender group at time t. On the right-hand side, the first and second terms capture the relative productivity of immigrants and the relative number of immigrants, respectively. The relative

[26]Even after the inclusion of our complete set of fixed effects (which drastically reduce the scope for omitted variable bias), it might still be the case that relative labor supplies are endogenous. For instance, a women-biased productivity shock in a specific cell could increase the cell-specific relative labor demand and relative wage of women, attracting relatively more women in those cells. In that case, the true value of the elasticity of substitution between men and women should be lower than our OLS estimates. We provide later an IV estimate of the elasticity of substitution which is lower than the OLS based elasticity.

productivity of immigrants can be captured by a set of fixed effects. As for the estimation of σ_F, we use model that includes all pair-wise interactions between education, experience, time and gender dummy variables. Each regression uses the inverse variance of the ratio of the log hourly wages as analytical weights (Borjas, Grogger, and Hanson, 2012). We cluster the standard errors at the education-experience-gender level.

– Insert Table 5 about here –

In Table 5, we report the estimated values of $-1/\sigma_I$.[27] The estimates are insignificant and, therefore, suggest that natives and immigrants of similar education and experience tend to be perfect substitutes in production.[28] This finding is in line with Edo and Toubal (2015) for France and Aydemir and Borjas (2007) for Canada and the United States. In order to simulate the impact of immigrant on wages, we assume σ_I to be infinity. However, given the mixed results regarding the degree of substitution between natives and immigrants reached in the literature we will also assume imperfect substitution between immigrants and natives.

Elasticities of substitution between education and experience groups. In order to simulate the wage effects of immigration for France, we also need to define the elasticity of substitution between the education and experience groups. We rely on Edo and Toubal (2015) who already estimated these parameters for France over the 1990-2010 period. We thus assume the elasticity of substitution between the two broad education groups $\sigma_{HL} = 4$ and between the medium and low educated groups $\sigma_L = 10$. We will test the robustness of our simulations by using an alternative set of parameters from Ottaviano and Peri (2012): $\sigma_{HL} = 2$ and $\sigma_L = 20$.

For the elasticity of substitution between experience groups, we assume that $\sigma_X = 7$. Our chosen value for σ_X is in line with Card and Lemieux (2001) who find an elasticity between 5 and 10, Ottaviano and Peri (2012) who find an elasticity around 6.25, Manacorda, Manning, and Wadsworth (2012) who find σ_X to be around 10 for the United Kingdom, and Edo and Toubal (2015) who find $\sigma_X = 7.5$ for France.

5 Empirical Results

We first provide an econometric analysis of the effect of immigration within group of workers with similar education and experience. We estimate the impact of the relative labor supply of female

[27] Notice that the number of observations is 504 as we take into account the gender dimension in our estimation. In Equation (11), we identify $-1/\sigma_I$ from changes that occur within gender-skill cells over time.

[28] Using population as a measure of labor supply (instead of employment) does not change our conclusions.

immigrants within skill group on the relative wages of female native workers. We then take into account the impact on native workers that are not competing in the same experience and education group by using a structural approach.

The within skill group estimations. The fact that within skill groups *(i)* men and women tend to be imperfect substitutes whereas *(ii)* immigrants and natives tend to be perfect substitutes has one important implication. An increase in the relative number of female immigrants within a skill group should affect negatively the relative wages of native women, widening the gender wage gap among native workers within that group. From Equation (10), we derive the following *reduced form* relative wage equation:[29]

$$log\left(\frac{w_{bjkt}^{Female}}{w_{bjkt}^{Male}}\right)^{N} \quad = \quad \beta \cdot log\left(\frac{m_{bjkt}^{Female}}{m_{bjkt}^{Male}}\right) + \delta_{bjk} + \delta_{kt} + \delta_{bjt} + \eta_{bjkt} \tag{12}$$

where m_{bjkt}^{Female} is one plus the ratio between the number of female immigrant workers and the number female native workers in the skill group (bj, k) and m_{bjkt}^{Male} is one plus the ratio between the number of male immigrant workers and the number of male native workers in the skill group (bj, k). m_{bjkt}^{Female} and m_{bjkt}^{Male} respectively measure the female and male immigrant supply shocks experienced by a particular skill group. The log of the ratio between m_{bjkt}^{Female} and m_{bjkt}^{Male} therefore captures the immigration contribution to changes in the relative labor supply of women.[30]

The use of a set time-varying fixed effects allows us to analyze how an immigration-induced change in the relative labor supply of women affects the gender wage gap of native workers with similar education and experience. When estimating Equation (12), we thus control for the fact

[29]By assuming a pre-immigration period and using Equation (10), the change of the relative wage of female natives N resulting from an exogenous inflow of male and female immigrants M is:

$$\begin{aligned}\triangle log\left(w_{bjkt}^{Female}/w_{bjkt}^{Male}\right)^{N} &= \triangle log\left(\theta_{S_F bjkt}/\theta_{S_M bjkt}\right)^{N} + \xi \cdot log\left[\frac{\left(Female_{bjkt}^{N}\left(1+n_{bjk}^{Female}\right) + Female_{bjkt}^{M}\right)/Female_{bjkt}^{N}}{\left(Male_{bjkt}^{N}\left(1+n_{bjk}^{Male}\right) + Male_{bjkt}^{M}\right)/Male_{bjkt}^{N}}\right] \\ &= \triangle log\left(\theta_{S_F bjkt}/\theta_{S_M bjkt}\right)^{N} + \xi \cdot log\left[\frac{1+n_{bjk}^{Female}+r_{bjkt}^{Female}}{1+n_{bjk}^{Male}+r_{bjkt}^{Male}}\right]\end{aligned}$$

$r_{bjkt}^{Female} = Female_{bjkt}^{M}/Female_{bjkt}^{N}$ and $r_{bjkt}^{Male} = Male_{bjkt}^{M}/Male_{bjkt}^{N}$: these measures of the relative supply of immigrants are cell specific and varies over time. n_{bjk}^{Female} and n_{bjk}^{Male} are the percent change in the relative number of female and male native workers. They are assumed to be cell specific and invariant over time.

[30]The derivation is parallel to the one of Borjas, Freeman, and Katz (1997, p. 41) in another context as they analyze the contribution of immigration to the change in the wage gap between skilled and unskilled workers in the U.S.

that the native gender wage gap may differ across education and experience groups over time by including all possible fixed effects. δ_{bjk} is a vector of fixed effects including education fixed effects (δ_{bj}), experience fixed effects (δ_k) and their interaction ($\delta_{bj} \times \delta_k$). This vector controls for all factors that are cell specific and invariant over time. The inclusions of δ_{kt} and δ_{bjt} control for the possibility that the gender wage gap may vary due to time-specific shocks, as well as for the fact that the impact of education and experience on relative wages may differ over time. As a result, β is identified using time-variation within education-experience cells. As noted in Ottaviano and Peri (2012), since we use relative wages and relative labor supplies at the skill-cell level, any (unobserved) specific shock at the skill-cell level that affects the labor market conditions of workers homogeneously should not bias our estimates.

The regression results are reported in Table 6. The dependent variable is the log ratio of native hourly wages. We weight each regression to account for differential precision in the measure of relative hourly wages across cells. As a weight, we thus follow Borjas, Grogger, and Hanson (2012) by using the inverse of the sampling variance of the dependent variable. We also cluster the standard errors by education-experience groups.

We investigate whether our results are robust to alternative definitions of the relative supply shock induced by female immigrant workers. We use:

- $log\left(Female\,Immigrants/Male\,Immigrants\right)_{bjkt}$. This ratio is very close to our baseline measure and facilitates the interpretation of the estimated coefficients.
- $log\left(Female\,Immigrants/All\,Immigrants\right)_{bjkt}$. This third measure of the supply shock induced by female/male immigrants uses the total number of immigrant workers (men and women) as a denominator.
- $\left(Female\,Immigrants/Male\,Immigrants\right)_{bjkt}$. This last measure does not take the log of the relative number of female immigrants.

In column 1 of Table 6, we use $log\left(m_{bjkt}^{Female}/m_{bjkt}^{Male}\right)$ as our baseline measure to capture the immigration-induced change in the relative supply of female workers. Columns 2, 3 and 4 respectively use our second, third and fourth alternative measures of the relative labor supply shock induced by female immigrants.

– Insert Table 6 about here –

The estimated coefficients are always negative and significant.[31] This finding is consistent

[31] In unreported regressions, we find that using relative monthly wages instead of relative hourly wages as dependent variable leads to more significant coefficients. For instance, the analogous specification to that reported in column 1 yields an estimated coefficient (t-student) of -0.02 (-3.71), implying a 1% significance level.

with the imperfect substitution elasticity between women and men. In columns 1 and 2, the estimated coefficients are similar. From column 2, our estimate implies that an increase of 10% in the relative number of female immigrants reduces the relative hourly wage of female natives with similar education and experience by 1%. As compared to column 2, column 4 shows that the within-cell impact of the relative supply of female immigrants on the native gender wage gap becomes more significant with no log transformation of the explanatory variable. Finally, the estimated coefficient from column 3 implies that a 10% increase in the share of female immigrant workers in a particular skill group reduces the native wage gap by 2%.[32]

The results reported in Table 6 point to a negative effect of the increase in the relative supply of female immigrants on the relative hourly wage of female native workers within the same education-experience group.

Endogenous labor supply. As relative labor supplies are likely to be endogenous, we implement a set of instrumental variables (IV) regressions. For instance, a women-biased productivity shock in a specific cell could increase the cell-specific relative labor demand and relative wage of women, attracting relatively more women in those cells. Thus, the OLS estimates from the within-cell regressions have to be interpreted as lower bounds of the true impact of an immigration-induced increase in the relative supply of women on their relative wage.

Our instrument is the relative number of single *immigrant* women within skill groups:

$$\left(single^{Female}_{Immigrant}/single^{Male}_{Immigrant}\right)_{bjkt}$$

This instrument is related to the relative labor force market participation of women. As suggested by the literature on the female labor market participation, the marital status should affect the reservation wages and not the market wages of women (Cain and Dooley 1976; Blackaby, Leslie, Murphy, and O'Leary 2002; Mulligan and Rubinstein 2008).[33] By affecting reservation wages, the marital status (being single or not) should thus be an important predictor of employment.[34] As a result, a change in the relative number of single immigrant women should affect the relative number of immigrant female workers. The lower-part of Table 7 supports this prediction: our first

[32]Column 3 indicates a stronger negative impact since the size of the denominator is larger than in column 2. In fact, the supply shock generated by a 10% increase in the relative number of female immigrants is larger in column 3 than in column 2.

[33]The labor force participation decision is generally based on a comparison of the market wage with the reservation wage. Individuals participate in the labor market when their offered wage exceeds their reservation wage (Killingsworth and Heckman, 1986; Heckman, 1993). Moreover, our instrument concerns strictly the population of immigrants and does not account for natives.

[34]This is shown empirically by Blackaby, Leslie, Murphy, and O'Leary (2002) for the United Kingdom.

stage estimated coefficients imply that our instrument is positively correlated with the relative number of immigrant female workers.[35] This positive correlation indicates that single individuals are more likely to participate in the labor market. This result is consistent with the findings of Mulligan and Rubinstein (2008).

The upper-part of Table 7 reports the IV estimates as well as their 95% confidence interval produced using the Conley, Hansen, and Rossi (2012) procedure.[36] As pointed out by Conley, Hansen, and Rossi (2012), it may be more credible to assume that the instrument does not fully satisfy the exclusion restriction. We relax the exclusion restriction and model the "plausible exogeneity" of our instrument by assuming that its estimated impact on the relative wage of native women lies in the interval $\gamma \in [-2\delta\,;\,2\delta]$.[37] By regressing the relative wage of native women on the relative number of single *immigrant* women (and all the covariates used in Equation 12), we find an estimated coefficient on the instrument equal to $\gamma = -0.01$ (with a corresponding t-student equal to -2.09). This result implies that $\delta = 0.005$. In order to estimate the 95% confidence interval of β (Equation 12), after allowing for some deviations from the IV exclusion restriction, we need to define the distribution of γ. We follow the suggestion by Conley, Hansen, and Rossi (2012, p. 266) and assume γ to be normally distributed with mean 0 and variance δ^2.

In all specifications of Table 7, the estimated coefficients are negative and of greater magnitude than our previous OLS estimates. This is consistent with an upward bias in the OLS estimates. More specifically, our estimate from column 2 implies that a 10% increase in the relative number of female immigrant workers lowers by 4% the relative wage of female native workers in the same education-experience group.

We also report the 95% confidence interval of our IV estimated coefficients by allowing for the exclusion restriction to not hold exactly. For each specification, the Conley, Hansen, and Rossi (2012) procedure indicates that none of the bounds contains zero. In our baseline specification (column 1), we find that under *perfect instrument* ($\gamma = 0$) the 95% confidence interval of $\hat{\beta}$ is [-0.08;-0.04], while it is [-0.11;-0.01] when allowing for some departures from *perfect instrument*. These results reinforce the evidence that an immigration-induced increase in the relative labor supply of women reduces the relative wage of native women. Therefore, the main result of the paper remains and is robust to deviations from the exclusion restriction imposed by the IV methodology.

[35] In addition, the F-test of exclusion is always higher than the lower bound of 10 suggested by the literature on weak instruments. We also implement the Conley, Hansen, and Rossi (2012) procedure to allow for deviations from the exclusion restriction.

[36] As before, we weight regressions by using the inverse of the sampling variance of the dependent variable.

[37] As proposed by Conley, Hansen, and Rossi (2012), we implement the "Local-to-Zero Approximation". The γ interval is also used by Conley, Hansen, and Rossi (2012, p. 266).

Results from the structural approach. Our structural estimations take explicitly into account the change in the gender composition of the immigrant workforce over the 1990-2010 period. This period is characterized by a tremendous increase in the share of female immigrants. Table 8 reports the simulated effects of immigration on the mean wage of native workers in the long-run along with the distributional effects of immigration by gender and education. The alternative parameter values of our simulations are reported in the upper-part of Table 8.[38] Table 11 in Appendix reports the confidence interval of the simulated effects. The effects are statistically significant at 1% level.

– Insert Table 8 about here –

As shown in Table 8, our simulated results indicate that immigration has contributed to a slight increase in the gender wage gap over a quite large parameter space. Our baseline long-run simulation indicates however that immigration has decreased the wage of French female natives by only 0.11% while it has increased the wage of male native by 0.06%, having therefore a weak effect on the gender wage gap. Given the results of the econometric exercise that exclusively analyses the within-cell effects of immigration, the structural approach seems to indicate a dampening effect of immigration when accounting for cross-group complementarities.[39]

In columns 1 to 4 of Table 8, we assume $\sigma_F = 12.5$.[40] In columns 1 and 2, we assume perfect substitution between natives and immigrants ($\sigma_I \to \infty$). We show that immigration has a negligible impact on the average wage of native workers in the long run irrespective of the degree of substitution between education groups. There is however a slight positive effect of immigration on the gender wage gap. In columns 3 and 4, we assume imperfect substitution between native and immigrants ($\sigma_I = 20$) as in Ottaviano and Peri (2012). In line with the results of Ottaviano and Peri (2012), we find positive simulated effects of immigration on the average wage of native workers, which are not depending on the the degree of substitution between education groups. Given the strong feminization of the immigrant workforce and imperfect substitutability between

[38] In order to compute the confidence intervals, we consider 500 draws of our set of substitution elasticities from the normal parameter distribution with the specified average and standard deviation (upper-part of Table 11). For each draw, we compute the total effects of immigration on native wages. This exercise thus produces 500 simulated effects from which we can compute the average effects of immigration as well as the corresponding standard errors and confidence interval.

[39] Notice however that the supply shocks in both analysis are different.

[40] This choice is derived from our previous OLS estimates presented in Table 4. As argued before, the true value of the elasticity of substitution between men and women should be lower than 12.5. We also implement an IV strategy using the number of single women relative to single men as an instrument. The IV coefficient based on our baseline specification (column 1, Table 4) is -0.13 (with a corresponding t-student equal to -2.17 ; the F-test from the first stage equal to 26.2). The IV estimate implies an elasticity of substitution between men and women with similar education and experience by around 7.7.

native and immigrants, these effects are mostly driven by a strong positive impact on male native workers while the effects on female native worker is negligible. The results of columns 3 and 4 confirm that the effect of immigration on the gender wage gap is not depending on the degree of substitution between native and immigrants.

In column 5, we increase the imperfect substitutability between women and men ($\sigma_F = 5$) and show that an immigration-induced increase in labor supply raises the wages of male workers and decreases the wages of female workers. As shown in column 6 and in line with our model, the effect of immigration on the wages of women and men is negligible and symmetric when we assume perfect substitutability between men and women ($\sigma_F \to \infty$) and between migrants and natives ($\sigma_I \to \infty$).

These different average effects (across columns) can be decomposed across education groups. In line with the fact that the share of high educated immigrants substantially increased over our period (as shown in Table 1), the negative wage effects of immigration are mainly driven by the wage losses experienced by the highly educated male and female native workers. More specifically, immigration has decreased their wages by between -0.4% and -1.0%. The asymmetric wage impact of immigration across education groups is reinforced in columns 2 and 4 when we assume $\sigma_{HL} = 2$ – *i.e.* a very low degree of substitution between the two broad education groups. In fact, when we assume $\sigma_{HL} = 2$, the wage effect induced by (a high educated) immigration is more concentrated among the highly educated natives, rather than diffused among all native workers. In addition, we find that both low and medium educated natives have experienced a slight improvement in their wage levels. Similar results are found in Edo and Toubal (2015).

In the lower-part of Table 8, we finally decompose the average wage effects of immigration across education groups by gender. As men and women are imperfect substitutes (columns 1 to 5), we find that immigration to France has increased the wage gap between male and female native workers with similar education. In fact, the imperfect substitutability between men and women induces weaker wage losses among the high educated male natives compared to female natives; as well as stronger wage gains among the medium and low educated native men.

Table 9 aims at comparing the mean effect of immigration on the gender wage gap under rigid labor market with the reference case of perfect labor markets. More specifically, this table reports the effects of immigration on the wages of female and male natives according to whether wages are assumed to be rigid (left-hand side) or perfectly flexible (right-hand side). We set $\sigma_{HL} = 4$ and $\sigma_L = 10$ (Edo and Toubal, 2015). For each scenario, we assume different degree of substitutability between men and women, as well as between immigrants and natives. For each simulation, we compute the difference between the wage changes of male and female native workers. A positive number implies that immigration has increased the gender wage gap.

– Insert Table 9 about here –

First, our results indicate that the negative effect of immigration on the gender wage gap is greater when σ_F is low and σ_I is high. Second, the impact of immigration on the relative wage of female natives is lower in rigid labor markets than in flexible labor markets. This finding points out the important role played by labor market rigidities in dampening the impact of immigration on the gender wage dispersion.

It is important to note that the two types of simulations (according to whether we assume rigid or flexible labor markets) can be interpreted as giving numerical bounds for the wage effects of immigration. Actually, it might be that the displacement effects due to immigration (inducing indirect wage effects) are not identical across all labor market cells, but rather concentrated in some skill-cells.[41] A reason for such asymmetric displacement effects may lie in the fact that some skill-cells are less affected by wage rigidities. As a result, the effect of immigration on the average wages of natives in France should be somewhere between a scenario where all wages are rigid and a scenario where wages are perfectly flexible.

6 Conclusion

France has experienced in the last two decades an important increase in the relative supply of female immigrant workers. We examine the impact of the feminization of the migration population on the wages of native female and male workers.

The impact of the feminization of the labor force on the gender wage gap depends on whether men and women are "poor" or "good" substitutes (Topel, 1997). If men and women are "poor" substitutes, the rising labor supply of immigrant women relative to immigrant men should therefore affect the gender wage gap. The degree of substitutability between men and women is therefore crucial for our study. We therefore develop a structural model to estimate the elasticity of substitution between men and women of similar education and experience. We find a significant degree of imperfect substitutability between men and women workers with similar education, experience and broad occupational categories. The literature in economics, sociology and psychology points to numerous drivers of this imperfect substitution (Maltz and Borker, 1982; Charles and Grusky, 2005; Croson and Gneezy, 2009).

In accordance with this imperfect substitution, we find that the immigration-induced increase in the relative supply of women in a given education-experience group reduces the relative wage

[41] In this regard, Edo (2015) finds that the negative effects of immigration on the employment of competing native workers are mostly concentrated within the medium and low educated segments of the French labor market.

of female natives in that group. The rise in the relative supply of women due to immigration therefore increased the wage gap between male and female native workers with similar education and experience.

Our structural model quantifies the distributional effects of female immigration by accounting for the cross-group effects induced by immigration. The long-run simulations reveal that the increasing feminization of the immigrant labor force has been detrimental for the relative wage of female natives, implying an increase in the gender wage gap. In particular, our baseline specification indicates that immigration has lowered the wage of female natives by 0.11% and increased the wage of male natives by about 0.06% over our period of interest (1990-2010).

Our results suggest that the gender composition of immigrants matters in determining their impact on the wage gap between male and female native workers.

References

Acemoglu, D., D. Autor, and D. Lyle (2004): "Women, war, and wages: The effect of female labor supply on the wage structure at midcentury," *Journal of Political Economy*, 112(3), 497–551.

Altonji, J., and R. Blank (1999): "Race and gender in the labor market," *Handbook of Labor Economics*, 3, 3143–3259.

Amuedo-Dorantes, C., and S. De La Rica (2011): "Complements or substitutes? Task specialization by gender and nativity in Spain," *Labour Economics*, 18(5), 697–707.

Angrist, J. D., and A. D. Kugler (2003): "Protective or counter-productive? labour market institutions and the effect of immigration on EU natives," *The Economic Journal*, 113(488), F302–F331.

Anker, R., et al. (1998): *Gender and jobs: Sex segregation of occupations in the world*. Cambridge Univ Press.

Autor, D. H., L. F. Katz, and M. S. Kearney (2008): "Trends in US wage inequality: Revising the revisionists," *The Review of Economics and Statistics*, 90(2), 300–323.

Aydemir, A., and G. J. Borjas (2007): "Cross-country variation in the impact of international migration: Canada, Mexico, and the United States," *Journal of the European Economic Association*, 5(4), 663–708.

Aydemir, A., and G. J. Borjas (2011): "Attenuation Bias in Measuring the Wage Impact of Immigration," *Journal of Labor Economics*, 29(1), 69–112.

Babeckỳ, J., P. Du Caju, T. Kosma, M. Lawless, J. Messina, and T. Rõõm (2010): "Downward Nominal and Real Wage Rigidity: Survey Evidence from European Firms*," *The Scandinavian Journal of Economics*, 112(4), 884–910.

Barone, G., and S. Mocetti (2011): "With a little help from abroad: the effect of low-skilled immigration on the female labour supply," *Labour Economics*, 18(5), 664–675.

Beauchemin, C., C. Borrel, and C. Régnard (2013): "Immigrants in France: a female majority.," *Population & Sociétés*, (502).

Blackaby, D. H., D. G. Leslie, P. D. Murphy, and N. C. O'Leary (2002): "White/ethnic minority earnings and employment differentials in Britain: evidence from the LFS," *Oxford Economic Papers*, 54(2), 270–297.

Blau, F., and L. Kahn (2000): "Gender Differences in Pay," *Journal of Economic Perspectives*, 14(4), 75–99.

Blau, F. D., M. A. Ferber, and A. E. Winkler (2002): *The economics of women, men, and work*, vol. 4. Prentice Hall Upper Saddle River, NJ.

Blau, F. D., and L. M. Kahn (2007): "Changes in the Labor Supply Behavior of Married Women: 1980-2000," *Journal of Labor Economics*, 25(3), 393–438.

Booth, A., and A. Leigh (2010): "Do employers discriminate by gender? A field experiment in female-dominated occupations," *Economics Letters*, 107(2), 236–238.

Borghans, L., B. Ter Weel, and B. A. Weinberg (2014): "People skills and the labor-market outcomes of underrepresented groups," *Industrial & Labor Relations Review*, 67(2), 287–334.

Borjas, G. J. (2003): "The labor demand curve is downward sloping: reexamining the impact of immigration on the labor market," *The quarterly journal of economics*, 118(4), 1335–1374.

Borjas, G. J. (2006): "Native internal migration and the labor market impact of immigration," *Journal of Human Resources*, 41(2), 221–258.

Borjas, G. J., R. B. Freeman, and L. F. Katz (1997): "How much do immigration and trade affect labor market outcomes?," *Brookings papers on economic activity*, 1997(1), 1–90.

Borjas, G. J., J. Grogger, and G. H. Hanson (2008): "Imperfect substitution between immigrants and natives: a reappraisal," Discussion paper, National Bureau of Economic Research.

——— (2011): "Substitution Between Immigrants, Natives, and Skill Groups," Discussion paper, National Bureau of Economic Research.

——— (2012): "Comment: On Estimating Elasticities Of Substition," *Journal of the European Economic Association*, 10(1), 198–210.

Borjas, G. J., and L. F. Katz (2007): "The evolution of the Mexican-born workforce in the United States," in *Mexican immigration to the United States*, pp. 13–56. University of Chicago Press.

Brücker, H., A. Hauptmann, E. J. Jahn, and R. Upward (2014): "Migration and imperfect labor markets: Theory and cross-country evidence from Denmark, Germany and the UK," *European Economic Review*, 66, 205–225.

Brücker, H., and E. J. Jahn (2011): "Migration and Wage-setting: Reassessing the Labor Market Effects of Migration," *The Scandinavian Journal of Economics*, 113(2), 286–317.

Cain, G. G., and M. D. Dooley (1976): "Estimation of a model of labor supply, fertility, and wages of married women," *Journal of Political Economy*, pp. S179–S199.

Card, D. (2009): "Immigration and Inequality," *The American Economic Review*, pp. 1–21.

——— (2012): "Comment: The elusive search for negative wage impacts of immigration," *Journal of the European Economic Association*, 10(1), 211–215.

Card, D., F. Kramarz, and T. Lemieux (1999): "Changes in the Relative Structure of Wages and Employment: A Comparison of the United States, Canada, and France," *The Canadian Journal of Economics*, 32(4), 843–877.

Card, D., and T. Lemieux (2001): "Can Falling Supply Explain the Rising Return to College for Younger Men? A Cohort-Based Analysis," *The Quarterly Journal of Economics*, 116(2), 705–746.

Charles, M., and D. B. Grusky (2005): *Occupational ghettos: The worldwide segregation of women and men*. Stanford University Press.

Cohen, D., A. Lefranc, and G. Saint-Paul (1997): "French unemployment: a transatlantic perspective," *Economic Policy*, 12(25), 265–292.

Collado, M. D., I. Iturbe-Ormaetxe, and G. Valera (2004): "Quantifying the impact of immigration on the Spanish welfare state," *International Tax and Public Finance*, 11(3), 335–353.

Conley, T. G., C. B. Hansen, and P. E. Rossi (2012): "Plausibly exogenous," *Review of Economics and Statistics*, 94(1), 260–272.

Cortes, P., and J. Pan (2015): "When Time Binds: Returns to Working Long Hours and the Gender Wage Gap among the Highly Skilled," *Mimeo*.

Cortes, P., and J. Tessada (2011): "Low-skilled immigration and the labor supply of highly skilled women," *American Economic Journal: Applied Economics*, 3(3), 88–123.

Croson, R., and U. Gneezy (2009): "Gender differences in preferences," *Journal of Economic literature*, pp. 448–474.

D'Amuri, F., G. I. Ottaviano, and G. Peri (2010): "The labor market impact of immigration in Western Germany in the 1990s," *European Economic Review*, 54(4), 550–570.

D'Amuri, F., and G. Peri (2014): "Immigration, jobs, and employment protection: Evidence from europe before and during the great recession," *Journal of the European Economic Association*, 12(2), 432–464.

Docquier, F., B. L. Lowell, and A. Marfouk (2009): "A gendered assessment of highly skilled emigration," *Population and Development Review*, 35(2), 297–321.

Docquier, F., A. Marfouk, S. Salomone, and K. Sekkat (2012): "Are skilled women more migratory than skilled men?," *World Development*, 40(2), 251–265.

Docquier, F., Ç. Ozden, and G. Peri (2013): "The Labour Market Effects of Immigration and Emigration in OECD Countries," *The Economic Journal*.

Dustmann, C., T. Frattini, and I. Preston (2007): *A study of migrant workers and the national minimum wage and enforcement issues that arise*. Low Pay Commission.

Dustmann, C., T. Frattini, and I. P. Preston (2013): "The effect of immigration along the distribution of wages," *The Review of Economic Studies*, 80(1), 145–173.

Eccles, J. S. (1994): "Understanding women's educational and occupational choices," *Psychology of women quarterly*, 18(4), 585–609.

Edo, A. (2015): "The impact of immigration on native wages and employment," *The BE Journal of Economic Analysis & Policy*, 15(3), 1151–1196.

——— (2016): "How do rigid labor markets absorb immigration? Evidence from France," *IZA Journal of Migration*, 5(1), 1–20.

Edo, A., and F. Toubal (2015): "Selective Immigration Policies and Wages Inequality," *Review of International Economics*, 23(1), 160–187.

Elsner, B. (2013): "Emigration and wages: The EU enlargement experiment," *Journal of International Economics*, 91(1), 154–163.

Evers, M., R. De Mooij, and D. Van Vuuren (2008): "The wage elasticity of labour supply: a synthesis of empirical estimates," *De Economist*, 156(1), 25–43.

Farré, L., L. González, and F. Ortega (2011): "Immigration, family responsibilities and the labor supply of skilled native women," *The BE Journal of Economic Analysis & Policy*, 11(1).

Felbermayr, G., W. Geis, and W. Kohler (2010): "Restrictive immigration policy in Germany: pains and gains foregone?," *Review of World Economics*, 146(1), 1–21.

Gerfin, M., and B. Kaiser (2010): "The Effects of Immigration on Wages: An Application of the Structural Skill-Cell Approach," *Swiss Journal of Economics and Statistics*, 146(4), 709–739.

Gilligan, C. (1982): "In a different voice: Psychological theory and women's development," *Cambridge, MA: Harvard*.

Glewwe, P. (1996): "The relevance of standard estimates of rates of return to schooling for education policy: A critical assessment," *Journal of Development economics*, 51(2), 267–290.

Glitz, A. (2012): "The labor market impact of immigration: A quasi-experiment exploiting immigrant location rules in Germany," *Journal of Labor Economics*, 30(1), 175–213.

Goldin, C. (2014): "A grand gender convergence: Its last chapter," *The American Economic Review*, 104(4), 1091–1119.

Goldin, C. D., and L. F. Katz (2009): *The race between education and technology*. Harvard University Press.

Heckman, J. J. (1979): "Sample Selection Bias as a Specification Error," *Econometrica*, 47(1).

Heckman, J. J. (1993): "What has been learned about labor supply in the past twenty years?," *The American Economic Review*, 83(2), 116–121.

Juhn, C., and D. I. Kim (1999): "The effects of rising female labor supply on male wages," *Journal of Labor Economics*, 17(1), 23–48.

Jurajda, Š. (2003): "Gender wage gap and segregation in enterprises and the public sector in late transition countries," *Journal of comparative Economics*, 31(2), 199–222.

Katz, L., and K. Murphy (1992): "Changes in relative wages, 1963–1987: supply and demand factors," *The Quarterly Journal of Economics*, 107(1), 35–78.

KILLINGSWORTH, M. R., AND J. J. HECKMAN (1986): "Female labor supply: A survey," *Handbook of labor economics*, 1, 103–204.

MALTZ, D. N., AND R. A. BORKER (1982): "A cultural approach to male-female miscommunication," *New York: Cambridge University Press*, pp. 168–185.

MANACORDA, M., A. MANNING, AND J. WADSWORTH (2012): "The Impact of Immigration on the Structure of Wages: Theory and Evidence from Britain," *Journal of the European Economic Association*.

MARCELLI, E. A., AND W. A. CORNELIUS (2001): "The changing profile of Mexican migrants to the United States: New evidence from California and Mexico," *Latin American Research Review*, pp. 105–131.

MARIANNE, B. (2011): "New perspectives on gender," *Handbook of labor economics*, 4, 1543–1590.

MINCER, J. A. (1974): *Schooling, Experience and Earnings*. New York: Columbia University Press.

MULLIGAN, C. B., AND Y. RUBINSTEIN (2008): "Selection, Investment, and Women's Relative Wages over Time," *The Quarterly Journal of Economics*, 123(3), pp. 1061–1110.

OECD (2007): "Perspectives des migrations internationales," Discussion paper, OECD Publishing.

OMELANIUK, I. (2005): "Gender, poverty reduction and migration," *World Bank*, pp. 1–18.

OTTAVIANO, G. I., AND G. PERI (2012): "Rethinking the effects of immigration on wages," *Journal of the European Economic Association*, 10, 152–197.

PELLIZZARI, M., M. PACCAGNELLA, AND G. DE GIORGI (2014): "Gender Complementarities in the Labor Market," *Research in Labor Economics*.

SIKORA, J., AND A. POKROPEK (2011): "Gendered career expectations of students: Perspectives from PISA 2006," Discussion paper, OECD Publishing.

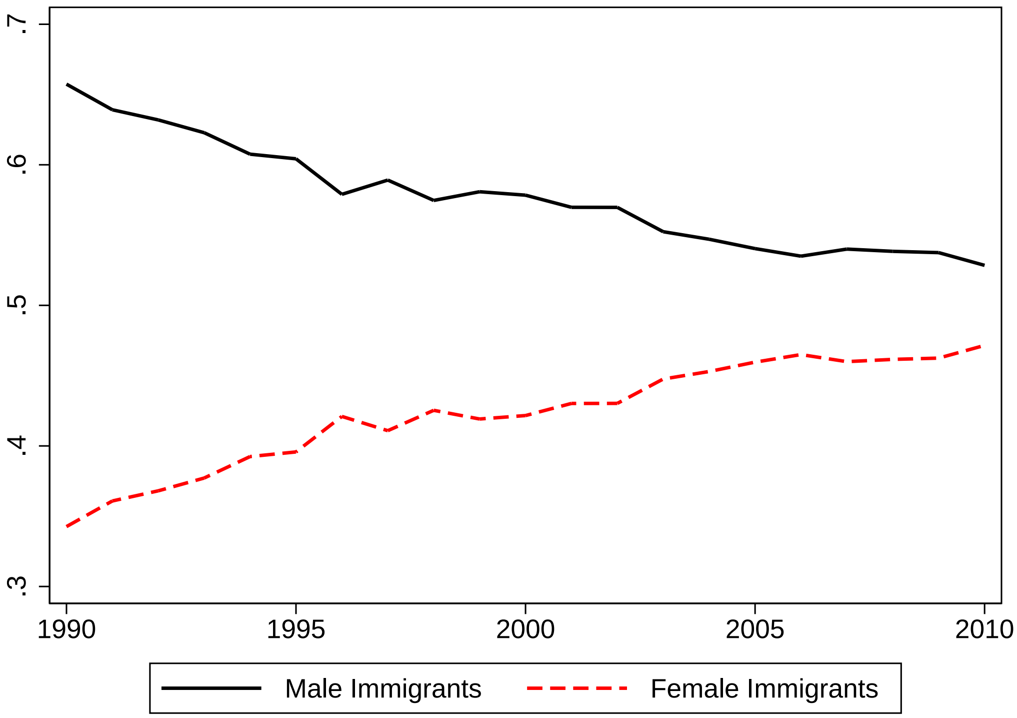
TOPEL, R. H. (1997): "Factor proportions and relative wages: the supply-side determinants of wage inequality," *The Journal of Economic Perspectives*, pp. 55–74.

UN (2013): "International Migration Report 2013," Discussion paper, Department of Economic and Social Affairs, Population Division.

Zlotnik, H. (1995): "The South-to-North migration of women," *International Migration Review*, pp. 229–254.

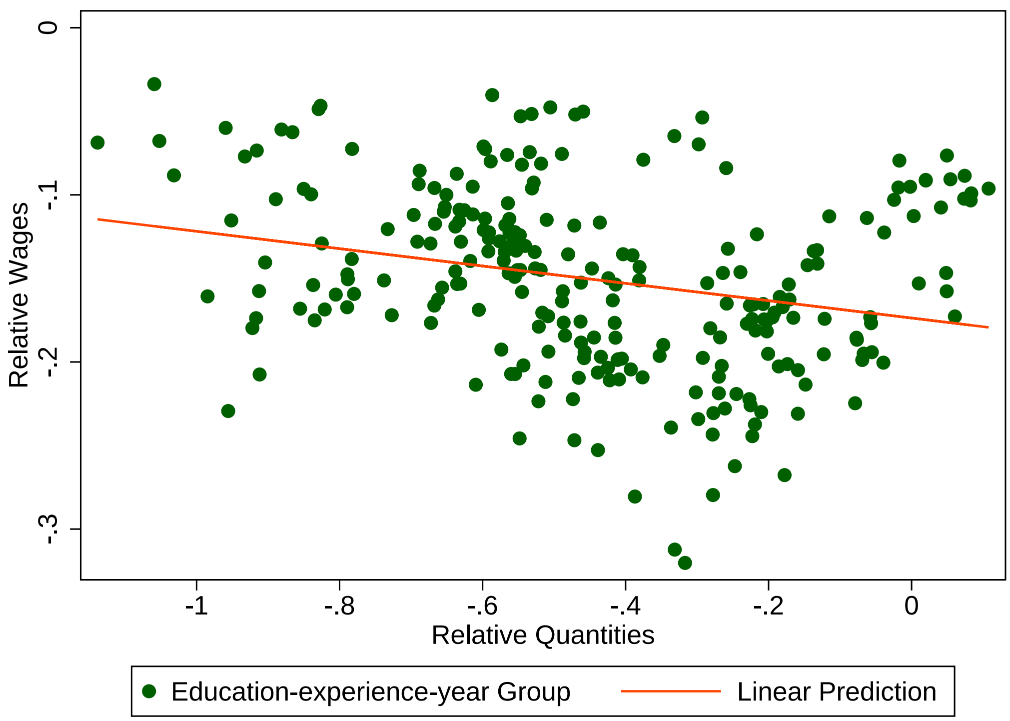
Figures

Figure 1: The Gender Distribution of Immigrants over Time



Notes. The Figure reports the shares of male and female immigrants in the immigrant labor force between 1990 and 2010. The population used to compute these shares includes all immigrants participating in the labor force aged from 16 to 64, not enrolled at school and having between 1 and 40 years of labor market experience. Self-employed people, workers in military occupations and the clergy are excluded from the sample.

Figure 2: Correlation between Relative Wage of Women and their Relative Labor Supply



Notes. The Figure provides the basic correlation between the relative wage of women and their relative labor supply by education-experience group in all years considered. The graphs give the log female-male hourly wage, $log(w^{F}_{bjkt}/w^{M}_{bjkt})$, on the vertical axis and the log female-male labor supply, $log\left(Female_{bjkt}/Male_{bjkt}\right)$, on the horizontal axis. We focus on the sample of full-time workers only. Each observation corresponds to an education-experience group in one of the considered year.

Figure 3: Evolution of the Gender Wage Gap and the Female Immigrant Supply Shift

Notes. The Figure reports the estimated gender wage gap for each year of our sample (left-axis) and the evolution of the relative number of female to male immigrants (right-axis).

Tables

Table 1: Educational Distribution by Gender and Origin in 1990 and 2010

	Males				Females			
	Natives		Immigrants		Natives		Immigrants	
	1990	2010	1990	2010	1990	2010	1990	2010
High Education	15.7 %	30.6 %	9.6 %	24.6 %	18.8 %	37.4 %	9.2 %	29.4 %
Medium Education	46.0 %	47.6 %	21.8 %	34.6 %	42.0 %	43.3 %	20.2 %	32.7 %
Low Education	38.2 %	21.8 %	68.7 %	40.8 %	39.2 %	19.4 %	70.7 %	37.9 %
Total	100 %	100 %	100 %	100 %	100 %	100 %	100 %	100 %

Table 2: Occupation by Gender and Origin (Average Share, 1990-2010)

	Males		Females	
	Natives	Immigrants	Natives	Immigrants
Professionals & Managers	17.7 %	11.9 %	10.4 %	8.7 %
Supervisors	25.1 %	14.1 %	25.1 %	12.9 %
Administrative Workers	14.7 %	13.2 %	52.7 %	55.3 %
Skilled-Manual Workers	25.6 %	33.9 %	3.9 %	6.0 %
Unskilled-Manual Workers	16.9 %	26.8 %	7.9 %	17.1 %
Total	100 %	100 %	100 %	100 %

Professionals and Managers: Manager of public sector; Professor and scientific profession; art, spectacles and information related professions; Administrative and commercial managers; Engineers and technical manager. **Supervisors:** School teachers, teachers and similar; Intermediate health professions and social work; Professionals in the public administration; Professionals in private companies; Technicians; Foremen, supervisors. **Administrative Workers:** Employees and agents of the public administration; Administrative employees; Commercial Workers Personal services. **Skilled-Manual Workers:** Skilled industrial workers; Skilled craft workers; Skilled handling, storage and transport. **Unskilled-Manuals Workers:** Drivers; Unskilled industrial workers; Unskilled craft workers; Laborers.

Table 3: Occupational Distribution by Gender, Origin and Education (1990-2010)

	Males		Females	
	Natives	Immigrants	Natives	Immigrants
	A. High Education			
Professionals & Managers	53.2 %	49.2 %	28.6 %	32.0 %
Supervisors	35.5 %	25.6 %	50.0 %	33.7 %
Administrative Workers	7.3 %	12.9 %	20.3 %	31.1 %
Skilled-Manual Workers	2.4 %	6.4 %	0.6 %	1.2 %
Unskilled-Manual Workers	1.6 %	5.8 %	0.5 %	1.9 %
Total	100 %	100 %	100 %	100 %
	B. Medium Education			
Professionals & Managers	8.1 %	5.6 %	3.9 %	4.3 %
Supervisors	25.9 %	19.1 %	18.8 %	13.1 %
Administrative Workers	16.1 %	15.2 %	66.7 %	64.9 %
Skilled-Manual Workers	34.1 %	39.3 %	4.1 %	6.6 %
Unskilled-Manual Workers	15.8 %	20.9 %	6.5 %	11.1 %
Total	100 %	100 %	100 %	100 %
	C. Low Education			
Professionals & Managers	4.9 %	1.5 %	1.7 %	0.7 %
Supervisors	15.3 %	7.0 %	9.3 %	3.5 %
Administrative Workers	18.4 %	12.3 %	64.2 %	60.7 %
Skilled-Manual Workers	30.4 %	41.2 %	7.1 %	7.8 %
Unskilled-Manual Workers	31.0 %	38.1 %	17.7 %	27.3 %
Total	100 %	100 %	100 %	100 %

Professionals and Managers: Manager of public sector; Professor and scientific profession; art, spectacles and information related professions; Administrative and commercial managers; Engineers and technical manager. **Supervisors:** School teachers, teachers and similar; Intermediate health professions and social work; Professionals in the public administration; Professionals in private companies; Technicians; Foremen, supervisors. **Administrative Workers:** Employees and agents of the public administration; Administrative employees; Commercial Workers Personal services. **Skilled-Manual Workers:** Skilled industrial workers; Skilled craft workers; Skilled handling, storage and transport. **Unskilled-Manuals Workers:** Drivers; Unskilled industrial workers; Unskilled craft workers; Laborers.

Table 4: Estimates of $-1/\sigma_F$, the Inverse Elasticity of Substitution between Men and Women

	Hourly Wage		
	Full-time Only	Full- and Part-time	Full-time Only
$-1/\hat{\sigma}_F$	**-0.07***	**-0.08****	**-0.06***
	(-2.03)	(-2.26)	(-1.70)
Education-Experience FE	Yes	Yes	-
Education-Time FE	Yes	Yes	-
Experience-Time FE	Yes	Yes	-
Occupation-Education-Experience FE	-	-	Yes
Occupation-Education-Time FE	-	-	Yes
Occupation-Experience-Time FE	-	-	Yes
Education-Experience-Time FE	-	-	Yes
Cluster	12	12	60
Observations	252	252	1,234

Key. ***, **, * denote statistical significance from zero at the 1%, 5%, 10% significance level. T-statistics are indicated in parentheses below the point estimate.
Notes. The dependent variable is the relative log mean wage of women. The explanatory variable is the log relative number of female employment in each cell. We use two different samples: one including part-time workers, one focusing on full-time workers only. All regressions include education, experience, and time fixed effects, as well as interactions between education and experience fixed effects, education and time fixed effects, and experience and time fixed effects. The standard errors are adjusted for clustering at the education and experience level.

Table 5: Estimates of $-1/\sigma_I$, the Inverse Elasticity of Substitution between Natives and Immigrants

	Hourly Wage	
	Full-time Only	Full- and Part-time
$-1/\hat{\sigma}_I$	**0.10**	**0.11**
	(1.11)	(0.97)
Education-Experience-Time FE	Yes	Yes
Gender-Education-Experience FE	Yes	Yes
Gender-Education-Time FE	Yes	Yes
Gender-Experience-Time FE	Yes	Yes
Cluster	24	24
Observations	504	504

Key. ***, **, * denote statistical significance from zero at the 1%, 5%, 10% significance level. T-statistics are indicated in parentheses below the point estimate.
Notes. The dependent variable is the relative log mean wage of immigrants. The explanatory variable is the log relative number of immigrant employment in each cell. We use two different samples: one including part-time workers, one focusing on full-time workers only. All regressions include a complete set of gender-education-experience fixed effects, gender-education-time fixed effects, gender-experience-time fixed effects, and education-experience-time fixed effects. The standard errors are adjusted for clustering at the gender-education-experience level.

Table 6: The Within-cell Effect of Immigrant Women on the Native Gender Wage Gap

	Explanatory Variable			
	Baseline	Log Female Immigrant Ratio	Log Female Immigrant Share	Female Immigrant Ratio
	$log\left(m_{bjkt}^{Female}/m_{bjkt}^{Male}\right)$	$log\left(M_{bjkt}^{Female}/M_{bjkt}^{Male}\right)$	$log\left(M_{bjkt}^{Female}/M_{bjkt}\right)$	$M_{bjkt}^{Female}/M_{bjkt}^{Male}$
Estimates	-0.01**	-0.01*	-0.02*	-0.03**
	(-3.03)	(-2.06)	(-1.88)	(-2.49)
Education-Experience FE	Yes	Yes	Yes	Yes
Education-Time FE	Yes	Yes	Yes	Yes
Experience-Time FE	Yes	Yes	Yes	Yes
Cluster	12	12	12	12
Observations	252	252	252	252

Key. ***, **, * denote statistical significance from zero at the 1%, 5%, 10% significance level. T-statistics are indicated in parentheses below the point estimate.
Notes. The dependent variable is the relative log mean hourly wage of native women. In column 1, our baseline explanatory variable is the log ratio between the relative number of female immigrants (as compared to female natives) and the relative number of male immigrants (as compared to male natives). In column 2, we take the log of the female/male immigrant ratio. In column 3, the explanatory variable is the log share of immigrant women. In column 4, we use the log relative number of immigrant women as compared to immigrant men. We focus our attention on full-time workers only. All regressions include education, experience, and time fixed effects, as well as interactions between education and experience fixed effects, education and time fixed effects, and experience and time fixed effects. The standard errors are adjusted for clustering at the education and experience level.

Table 7: The Within-cell Effect of Immigrant Women on the Native Gender Wage Gap

		Explanatory Variable			
		Baseline	Log Female Immigrant Ratio	Log Female Immigrant Share	Female Immigrant Ratio
		$log\left(m_{bjkt}^{Female}/m_{bjkt}^{Male}\right)$	$log\left(M_{bjkt}^{Female}/M_{bjkt}^{Male}\right)$	$log\left(M_{bjkt}^{Female}/M_{bjkt}\right)$	$M_{bjkt}^{Female}/M_{bjkt}^{Male}$
IV Estimates		-0.06***	-0.04***	-0.06***	-0.07***
		(-6.54)	(-5.17)	(-5.43)	(-4.46)
95% CI of IV Estimates		[-0.08 ; -0.04]	[-0.05 ; -0.02]	[-0.08 ; -0.04]	[-0.10 ; -0.04]
95% CI using Conley et al. (2012)		[-0.11 ; -0.01]	[-0.07 ; -0.01]	[-0.11 ; -0.01]	[-0.13 ; -0.01]
First Stage:	Instrument	0.23***	0.36***	0.23***	0.19***
	T-statistic	(3.89)	(6.53)	(5.88)	(8.13)
	F-statistic	15.76	42.72	34.75	66.41
Education-Experience FE		Yes	Yes	Yes	Yes
Education-Time FE		Yes	Yes	Yes	Yes
Experience-Time FE		Yes	Yes	Yes	Yes
Cluster		12	12	12	12
Observations		252	252	252	252

Key. ***, **, * denote statistical significance from zero at the 1%, 5%, 10% significance level. T-statistics are indicated in parentheses below the point estimate.
Notes. The dependent variable is the relative log mean hourly wage of native women. We use the same explanatory variables as in Table 6. We focus our attention on full-time workers only. All regressions include education, experience, and time fixed effects, as well as interactions between education and experience fixed effects, education and time fixed effects, and experience and time fixed effects. As instrument, we use the number of single immigrant women relative to single immigrant men. For each explanatory variable, we provide the IV estimated coefficient and its 95% confidence interval using the procedure by Conley, Hansen, and Rossi (2012) by assuming that $\gamma\epsilon[-2\delta; 2\delta]$ and $\gamma \sim \mathcal{N}\left(0, \delta^2\right)$. The standard errors are adjusted for clustering at the education and experience level.

Table 8: The Long-term Effects of Immigration on Native Wages by Gender and Education

	(1)	(2)	(3)	(4)	(5)	(6)
	Set of Parameter Values					
σ_{HL}	4	2	4	2	4	4
σ_L	10	20	10	20	10	10
σ_X	7	7	7	7	7	7
σ_F	12.5	12.5	12.5	12.5	5	∞
σ_I	∞	∞	20	20	∞	∞
	Percentage Change of Native Wage due to Immigration					
Average Long-term Effects	-0.01	-0.01	0.08	0.08	-0.01	-0.01
Male	0.06	0.06	0.10	0.11	0.16	-0.01
Female	-0.11	-0.13	0.04	0.03	-0.28	0.00
Highly Educated	-0.57	-1.14	-0.37	-0.95	-0.57	-0.57
Male	-0.49	-1.06	-0.34	-0.92	-0.36	-0.57
Female	-0.69	-1.26	-0.42	-0.99	-0.88	-0.56
Medium Educated	0.10	0.35	0.20	0.44	0.10	0.10
Male	0.17	0.42	0.22	0.46	0.30	0.08
Female	-0.01	0.23	0.16	0.41	-0.25	0.14
Low Educated	0.33	0.46	0.31	0.44	0.33	0.33
Male	0.35	0.48	0.32	0.45	0.37	0.33
Female	0.30	0.43	0.30	0.43	0.25	0.33

Notes. The table reports the simulated effects of immigration on the wages of native workers by education and gender. Each number stands for the percentage wage changes due to immigrants for natives. All simulations assume $\sigma_X = 7$. While columns 1 to 4 assume $\sigma_F = 12.5$, columns 5 and 6 assume alternative values for σ_F. We test the robustness of our results by using different substitution elasticity values between education groups, and between natives and immigrants. The total wage effect is computed as the sum of direct effects due to immigration and indirect effects due to employment responses.

Table 9: The Long-run Effects of Immigration by Gender under Rigid/Perfect Labor Market

	(1)	(2)	(3)	(4)	(5)	(6)
	Set of Parameter Values					
σ_{HL}	4	4	4	4	4	4
σ_L	10	10	10	10	10	10
σ_X	7	7	7	7	7	7
σ_F	5	12.5	12.5	5	12.5	12.5
σ_I	∞	∞	20	∞	∞	20
	Rigid Labor Market			**Perfect Labor Market**		
Average Wage Effect	-0.01	-0.01	0.08	-0.02	-0.02	0.26
Male (a)	0.16	0.06	0.10	0.32	0.12	0.31
Female (b)	-0.28	-0.11	0.04	-0.61	-0.25	0.17
Differences ($a - b$)	**0.44**	**0.17**	**0.06**	**0.93**	**0.37**	**0.18**

Notes. The table reports the long-run simulated effects of immigration on native wages by gender according to whether wages are rigid or perfectly flexible. Each number stands for the percentage wage changes due to immigrants for natives. We use $\sigma_{HL} = 4$, $\sigma_L = 10$, $\sigma_X = 7$ and different values for σ_I and σ_F. On the left-hand side, the total wage effect is computed as the sum of direct effects due to immigration and indirect effects due to employment responses.

Appendix

Conditional gender wage gap and the shift in the structure of the immigrant workforce. Our first step is to provide yearly estimates of the wage gap using the available microeconomic dataset. As wages are observed at a microeconomic level for a non-random subsample of individuals, namely the ones working, our empirical model controls for selection bias using the Heckman two-stage estimation procedure (Heckman, 1979) which provides us with selection-corrected estimates. We estimate the gender wage gap at the individual level by estimating the following wage equation :

$$w_{ijk} = \beta_0 + \beta_1 S_i + \beta_3 \lambda_i + \delta_j + \delta_k + \delta_j \times \delta_k + \varepsilon_{ijk} \tag{13}$$

w_{ijk} is the real hourly wage of individual i (in logarithm) with education j and experience k. As in the paper, we use three education groups (tertiary, secondary and primary education) and four experience groups of five years (1-10 ; 10-20 ; 20-30 ; 30-40). The gender status of an individual is captured by the gender dummy S_i which is equal to one when the worker is a male. The inverse mills ratio is $\lambda_i = \frac{\phi(Z\gamma)}{\Phi(Z\gamma)}$; Z is a set of regressors including the vector of explanatory variables of the participation equation and y is the first stage dependent variable. We derive the inverse mills ratio from our first-stage regression, where we estimate a selection equation by maximum likelihood as an independent probit model to determine the decision to enter the labor market. As selection variables, we rely on Glewwe (1996) and Mulligan and Rubinstein (2008) and use the marital status and its interaction with the number of children aged less than 6 years old. In the first- and second-stage regressions, we include vectors of fixed effects for education δ_j, experience δ_k as well as their interaction $\delta_j \times \delta_k$.[42] The linear fixed effects control for differences in labor market outcomes across education groups and experience groups, whereas the interaction controls for the fact that the impact of education on labor market outcomes may vary across experience groups. We therefore exploit the within-cell variation to identify the effects of gender on the probability to be employed and wage. We always correct the standard errors for heteroscedasticity.

Table 10 reports the estimated coefficient on the gender dummy for the first and second stage regressions for the years 1990, 2000 and 2010. In order to estimate the gender wage gap, we focus on the sample of full-time workers. The first-stage estimates always indicate that women are less likely to be in employment relative to men with similar education and experience characteristics.

[42] The results are also robust to the introduction of occupation fixed effects and their interactions with the education and experience dummy variables.

Table 10: Gender Differences in Wages and Employment

	1990		2000		2010	
	$P(N=1)$	Wage	$P(N=1)$	Wage	$P(N=1)$	Wage
Sex dummy	0.08***	0.07***	0.08***	0.10***	0.04***	0.12***
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Mills Ratio	-	-0.62***	-	-0.37***	-	-0.43***
		(0.04)		(0.04)		(0.04)
Selection Variables	Yes	-	Yes	-	Yes	-
Education-Experience FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	46,643	41,332	50,117	43,713	46,343	41,394

Notes. ***, **, * denote statistical significance from zero at the 1%, 5%, 10% significance level. All regressions include education and experience fixed effects, as well as their interaction. As selection variables, we use the marital status and its interaction with the number of children aged less than 6 years old. Robust standard errors are in parentheses below the point estimate.

Regarding the second-stage regressions (*i.e*, the wage regression), we find that the inverse Mills ratio is always negative and significant, highlighting a selectivity into employment. The negative selectivity term suggests that if unemployed people were to find a job, they would have higher earnings as compared to individuals with similar characteristics already in jobs. This result is consistent with the notion that unemployed people tend to have higher reservation wages than workers.

The estimated coefficients on the gender dummy has almost doubled over the period of estimation from around 0.06 to 0.12 – *i.e.*, within the same education-experience group, female workers earn 6 to 13 percent less than men. At the skill-cell level, we find that the immigrant wage gap is generally higher than the native wage gap (between 7 to 22 percent for immigrants and 6 to 13 percent for natives).

Table 11: The Confidence Intervals of the Simulated Effects of Immigration on Native Wages

	(1)	(2)	(3)	(4)	(5)	(6)
	Set of Parameter Values					
$-1/\sigma_{HL}$	-0.25	**-0.50**	-0.25	**-0.50**	-0.25	-0.25
	(0.03)	(0.06)	(0.03)	(0.06)	(0.03)	(0.03)
$-1/\sigma_{L}$	-0.10	**-0.05**	-0.10	**-0.05**	-0.10	-0.10
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
$-1/\sigma_{X}$	-0.14	-0.14	-0.14	-0.14	-0.14	-0.14
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
$-1/\sigma_{F}$	-0.08	-0.08	-0.08	-0.08	**-0.20**	**0.00**
	(0.03)	(0.03)	(0.03)	(0.03)	(0.05)	(0.00)
$-1/\sigma_{I}$	0.00	0.00	**-0.05**	**-0.05**	0.00	0.00
	(0.00)	(0.00)	(0.03)	(0.01)	(0.00)	(0.00)
Average Wage Effect	[-0.0060;-0.0058]	[-0.0079;-0.0077]	[0.078;-0.0813]	[0.077;-0.079]	[-0.0069;-0.0067]	[-0.0053;-0.0052]
Male	[0.055;0.061]	[0.061;0.067]	[0.103;0.109]	[0.109;0.115]	[0.148;0.158]	[-0.008;-0.0076]
Female	[-0.119;-0.109]	[-0.134;-0.124]	[0.030;0.042]	[0.015;0.028]	[-0.285;-0.268]	[-0.001;-0.0007]
Highly Educated	[-0,578;-0,567]	[-1,147;-1,125]	[-0,381;-0,369]	[-0,950;-0,927]	[-0,579;-0,568]	[-0,578;-0,567]
Male	[-0,503;-0,486]	[-1,074;-1,042]	[-0,357;-0,339]	[-0,927;-0,895]	[-0,381;-0,362]	[-0,584;-0,568]
Female	[-0,702;-0,682]	[-1,272;-1,239]	[-0,428;-0,406]	[-0,998;-0,963]	[-0,894;-0,869]	[-0,575;-0,559]
Medium Educated	[0,103;0,108]	[0,347;0,355]	[0,196;0,202]	[0,440;0,448]	[0,102;0,107]	[0,104;0,108]
Male	[0,167;0,177]	[0,410;0,424]	[0,212;0,223]	[0,456;0,470]	[0,294;0,309]	[0,083;0,089]
Female	[-0,018;-0,003]	[0,226;0,244]	[0,158;0,175]	[0,402;0,421]	[-0,251;-0,228]	[0,137;0,145]
Low Educated	[0,327;0,333]	[0,457;0,465]	[0,309;0,315]	[0,438;0,447]	[0,326;0,332]	[0,328;0,335]
Male	[0,343;0,352]	[0,472;0,484]	[0,313;0,322]	[0,442;0,455]	[0,368;0,377]	[0,327;0,335]
Female	[0,296;0,305]	[0,425;0,438]	[0,297;0,305]	[0,425;0,438]	[0,248;0,258]	[0,328;0,336]

Notes. The table reports the 99% confidence interval of the simulated long-run wage effects of immigration for the same specifications as in Table 8. We use the parameter values listed in the upper-part of the table (std. errors in parentheses), and for each specification, we generate 500 extractions of the parameters from a normal distribution. We are thus able to simulate the wage effect of immigration 500 times and obtain the mean simulated effect as well as the 99% confidence interval.