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Skill Premium, Labor Supply, and Changes in the Structure of Wages in Latin America*

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July 12, 2018

Abstract

After a decade of increasing wage inequality, this paper documents a sharp compression in the distribution of wages in Argentina and Chile during the 2000s. In Brazil, wage inequality has steadily declined since the early 1990s. Counterfactual exercises show that the evolutions of the schooling and experience premiums are key determinants of the decline in inequality. The 2000s witnessed a rapid decline in the schooling and experience premiums, at the same time as the working population was aging and becoming more educated. To understand these changes, the paper develops a model of imperfect substitution across experience and education groups and estimates the relevant elasticities of substitution. Changes in labor supply contributed significantly to the decline of the experience and education premiums, but are not enough to account fully for the observed changes. The demand for experience shifted in favor of younger workers, while the relative demand for college graduates declined during the 2000s.

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Abstract

After a decade of increasing wage inequality, this paper documents a sharp compression in the distribution of wages in Argentina and Chile during the 2000s. In Brazil, wage inequality has steadily declined since the early 1990s. Counterfactual exercises show that the evolutions of the schooling and experience premiums are key determinants of the decline in inequality. The 2000s witnessed a rapid decline in the schooling and experience premiums, at the same time as the working population was aging and becoming more educated. To understand these changes, the paper develops a model of imperfect substitution across experience and education groups and estimates the relevant elasticities of substitution. Changes in labor supply contributed significantly to the decline of the experience and education premiums, but are not enough to account fully for the observed changes. The demand for experience shifted in favor of younger workers, while the relative demand for college graduates declined during the 2000s.

1. Introduction

Inequality declined sharply in Latin American countries after the turn of the century, a contrast with the region's history and global trends (Ferreira et al., 2008; Kahhat, 2010; López-Calva and Lustig, 2010; Gasparini and Lustig, 2011; Gasparini et al., 2011a; Levy and Schady, 2013; Lustig et al., 2013; Messina and Silva, 2018). Redistribution through progressive fiscal policy, the emergence of conditional cash transfer programs for the poor, and changes in household demographics played a role in this transition. However, a broad conclusion of previous literature is that the key contribution to inequality reduction was the decline in earnings inequality (Lopez-Calva and Lustig, 2010; Azevedo et al., 2013). Earnings inequality declined in 16 of the 17 countries in Latin America for which consistent statistics can be calculated (Messina and Silva, 2018), although the intensity and turning points diverged across countries. For example, after a decade of stagnant or slowly increasing inequality, the 90th/10th interquantile range of the labor earnings distribution declined by 20 percent in Argentina and 28 percent in Chile between 2000 and 2013. In Brazil, where earnings inequality started to fall as early as 1990, the reduction has been a remarkable 46 percent since 2000.

At the time these sharp changes in the wage structure were taking place, the region was registering a rapid transformation in the composition of employment. The shares of women, highly educated, and older workers increased in most countries. Between 1990 and 2013, the share of college-educated workers increased from 19.9 to 29.1 percent in Argentina, virtually doubled in Chile (14.8 to 27.3 percent), and almost tripled in Brazil (7.8 to 19.5 percent). The average worker's age increased by more than a year in Argentina (37.3 to 39.0), by three years in Brazil (34.1 to 37.4), and by more than four years in Chile (35.8 to 40.4). The share of female employment increased in all countries, with Brazil and Chile ahead of the pack: an increase of more than 7 and 10 percentage points, respectively. Changes in the composition of the labor force can affect wage inequality because different types of workers have different levels of within-group wage dispersion (Lemieux, 2006). They can also affect inequality by changing between-group differences in pay through their effects on the education, gender, and experience premiums, as the canonical supply-demand framework of the labor market suggests.

This paper investigates how these changes in the demographic and skill structure of the labor force influenced the evolution of the distribution of earnings in Argentina, Chile, and Brazil during the past 25 years, using household survey data. Our analysis has two parts. The first part disentangles the contribution of pure compositional changes from changes in the structure of pay in the evolution of inequality. Following the work of Firpo et al. (2007, 2009), we construct counterfactual wage distributions that decompose the observed changes in inequality measures into price and composition effects. The analysis distinguishes between overall wage inequality (measured by the log interquantile 90/10 earnings ratio), lower tail inequality (measured by the log 50th/10th interquantile ratio), and upper tail inequality (measured by the log 90th/10th interquantile ratio). This is important because the trends have been different, an oft-neglected fact in discussions about inequality in the Latin America region. While the most significant portion of the reduction in earnings inequality in Brazil was the result of a decline in the 50th/10th interquantile range (-31 percent since 2000), the reduction in inequality in Argentina was mostly driven by a fall of upper tail inequality (-23 percent fall of the 90/50 log earnings ratio since 2000). In Chile, inequality fell symmetrically at the lower and upper parts of the earnings distribution (-15 percent since 2000).

In line with previous literature (López-Calva and Lustig, 2010; Gasparini and Lustig, 2011; Levy and Schady, 2013; Lustig et al., 2013), we find that the decline of the education premium is a key determinant of the observed wage inequality decline. However, the role of the decline of the education premium was very different in the upper and lower tails of the distribution. The decline of the returns to schooling unambiguously contributed to the decline of inequality at the bottom of the distribution, but, except for Chile, it had no effect on the reduction of upper tail inequality. This in turn reflects the diverging evolution of the high school and college premiums. While the high school/primary wage gap fundamentally fell throughout the period (with mild increases in the early 1990s in Chile and in the late 1990s in Argentina), the college/high school wage gap increased strongly during the 1990s and then rapidly declined during the 2000s. In Brazil and Chile, the college/high school wage gap in 2013 was essentially at the same level as it was in 1990. In Argentina, it was 14 log points below 1995, which is the first year we observe for this country (Figure 2).

Against the dominating role of falling schooling premiums, pure compositional changes related to the increase in educational attainment were inequality enhancing, thus contributing to increasing inequality in the early 1990s in Argentina and Chile but working against the recent inequality decline. This may be due to within-group differences in pay, as highlighted by Lemieux

(2006) in the United States, or it may reflect a phenomenon previously labeled as the paradox of progress (Bourguignon et al., 2005), by which increases in educational attainment can be inequality-increasing due to convexity of the returns to education.

The evolution of the experience premium has received considerably less attention in the literature that studies the determinants of wage dispersion in Latin America, which has focused on the returns to education.[1] However, the experience premium has a prominent role. It declined in the three countries and for all education groups (Figure 2), although there is significant heterogeneity. In Argentina, the wage gap of workers with 20-29 years of experience versus those with 0-9 years of experience (log 20-29/0-9 wage gap) declined sharply among workers with college (a decline of 26 log points) and high school education (-18 log points), remaining basically unchanged among workers with at most primary education (including high school dropouts). As a result, the Oaxaca-Blinder decompositions discussed in the paper show that the decline of the experience premium explains virtually the full compression of upper tail inequality observed in 1995-2013 in Argentina, while it has no statistically significant effect on the evolution of lower tail inequality. In Brazil and Chile, the fall of the experience premium is important across all schooling groups, and somewhat largest among workers with high school education. The log 20-29/0-9 wage gap of high school graduates fell by as much as 47 and 35 log points in Brazil and Chile, respectively. The widespread reduction of the experience premium in Brazil and Chile is reflected in the Oaxaca-Blinder decompositions for these two countries, which show important equalizing effects of changes in the returns to experience in the upper and lower tails of the distribution.

The second part of the analysis studies the role of the aforementioned labor supply changes in the evolution of the education and experience premiums. Have declines in those premiums been driven by the increasing educational attainment and aging of the labor force? To answer this question,

[1] The influential book by López-Calva and Lustig (2010) offers a comprehensive study of the decline in inequality in Argentina, Brazil, Mexico, and Peru. On the role of wages, a key conclusion is that the decline of wage inequality was fundamentally driven by changes in educational attainment. More recent works for Brazil reach similar conclusions. Wang et al. (2016) attribute a prominent role to falling returns to education in their analysis of the evolution of wage dispersion, although in their decompositions changes in the structure of pay associated with age present a very strong wage-compressing effect, which is not discussed in the paper. Jaume (2018) shows that increasing education has contributed to declining inequality through occupational downgrading, which is the focus in the paper. A similar mechanism could be at work regarding work experience. However, this is not discussed in the paper.

our modeling approach is closely linked to that of Katz and Murphy (1992), who use a CES production function with two education levels to link relative labor supply to the evolution of the education premium in the United States. This approach was extended by Card and Lemieux (2001) to introduce imperfect substitution across experience groups using data from Canada, the United Kingdom, and the United States. Manacorda et al. (2010) further extend the approach of Card and Lemieux (2001) to consider two types of workers among the unskilled group, high school dropouts and workers with at most secondary education completed, and estimate the model with data for Latin America.

We extend the model in Manacorda et al. (2010) and Card and Lemieux (2001) in two significant ways. First, we allow for differential demand trends across experience groups, while Manacorda et al. (2010) and Card and Lemieux (2001) rule out experience-biased demand changes. Ruling out experience-biased demand change imposes that all movements in the experience premium must be captured by changes in relative supply, a feature that is rejected in our data. Second, our model allows for different degrees of imperfect substitution across experience groups, depending on their level of schooling. We show that this distinction is empirically relevant. The estimated elasticity of substitution across experience groups is higher among college graduates than among unskilled workers. In contrast, Manacorda et al. (2010) find that workers of different age groups are perfect substitutes in production. Beyond these two extensions, and unlike previous papers using similar frameworks as ours, we take into account the uncertainty introduced by generated regressors in the estimation, by bootstrapping the standard errors. This turns out to be important for inference in this context. The bootstrapped standard errors are up to two times larger than the conventional heteroskedasticity-robust ones.

We find that the secular decline of the experience premium is consistent with the observed movements in labor supply in the three countries. However, our estimates of the elasticity of substitution suggest that changes in labor supply fall short of fully accounting for the observed dynamics. The estimated elasticity of substitution across experience groups within unskilled workers is 3.6. Using this estimate and the observed changes in labor supply, we conclude that labor supply accounts for about 50 percent of the observed reduction in the experience premium. For workers with college education, we estimate a higher elasticity of substitution across experience groups (about 5.5 in our preferred specification), and consequently labor supply changes account for at most a third of the decline. This finding indicates that during the past few decades demand forces have moved against experienced workers in Latin America. To our knowledge, this paper is the first attempt to account for the rapid decline of the experience premium in Latin American countries.

Changes in labor supply do a better job at capturing the dynamics of the high school/primary wage gap. Our estimate of the elasticity of substitution is 2.3. In Brazil, the sharp increase in the supply of high school graduates predicts a massive decline of the high school premium (-68.4 log points), which is very close to the observed reduction of the log high school/primary wage gap (-71.9 log points). In Argentina and Chile, the increase in the relative supply of high school-educated workers overshoots the observed decline in the high school premium. Thus, demand forces favoring high school-educated workers over dropouts have sustained the high school premium in the two countries.

The picture for college-educated workers is slightly different. With an estimated elastictiy of substitution of 1.25, we show that the rising supply of college-educated workers has pushed the college/unskilled wage gap downward over the past 25 years. Over the whole period, the evolution of labor supply slightly over-predicts the observed decline. As before, the fit of the labor supply model for Brazil is quite good. However, for Argentina and Chile, labor supply changes miss important dynamics in the college premium, which trended upward until the early 2000s and declined thereafter. This hump-shaped evolution of the college premium is captured by relative demand trends, which favored college-educated workers during the 1990s, but this started declining at the start of the new millennium. The implication is that demand-side trends attenuated labor supply forces toward a declining college premium during the 1990s, but accentuated the decline after 2003.

The analysis of the evolution of the schooling premium relates to the study by Galiani et al. (2017), on the role of education upgrading for the college and high school premiums in 16 Latin American countries. Building on Katz and Murphy (1992), Galiani et al. (2017) evaluate the role of labor supply changes in the evolution of the high school/primary and college/unskilled wage gaps for different values of the elasticity of substitution, which are obtained from the literature. Demand changes are the residual. Thus, there are two key differences between their approach and ours. First, we estimate the elasticities of substitution from the data, allowing for different values for college/unskilled and high school/primary educated workers. Second, Galiani et al. (2017) do not study the role of experience, and hence assume perfect substitution across experience groups. We instead build aggregate labor supply series of workers with different schooling levels, taking into account imperfect substitution and changes in relative efficiency across experience groups within each schooling category. Despite these noteworthy methodological differences, our findings are broadly consistent with theirs for a larger set of Latin American countries. They find that the labor supply model over-predicts the decline of the high school/dropout wage gap, as evidenced by an upward trend of residual demand (see Figure 4.3 in Galiani et al. (2017)). As we report for Argentina and Chile, their residual demand index for the college/unskilled wage gap follows an inverse U-shaped pattern, peaking around 2003.

The rest of the paper is organized as follows. Section 2 discusses the data and main stylized facts, reviewing the evolution of inequality and socio-demographic changes in Argentina, Brazil, and Chile. Section 3 shows how the evolution of inequality was affected by changes in composition and wage structure associated with education, experience, and gender. Section 4 discusses skill supply and demand and the evolution of relative returns. Section 5 develops a stylized model of supply and demand based on the descriptive trends in the data, and Section 6 provides estimates of the key parameters of the model and discusses its implications for the evolution of the skill premium. Section 7 concludes.

2. Data and Inequality Trends

We use household surveys for Argentina, Brazil, and Chile for the analysis. All the surveys include information about general characteristics of the workers (e.g., gender, age, education) and their jobs (type of contract, labor earnings, hours worked). With the exception of Argentina, where information is restricted to urban areas,[2] the other surveys are nationally representative. Earnings refer in the three surveys to total monetary payments from labor in a reference period. Labor earnings are divided by actual hours worked during the same period to obtain hourly earnings. The series are converted into real terms using the Consumer Price Index (CPI).[3] We restrict the sample to individuals between the ages of 16 and 65 and only use the earnings of full-time

[2] Urban areas accounted for almost 90 percent of the total population in Argentina in 2013.

[3] The official CPI is used in Brazil and Chile. Due to inconsistencies found in the official series in Argentina (see Cavallo (2013)), we use the information from PriceStats (http://www.statestreet.com/ideas/pricestats.html) instead. Because the paper focuses on inequality, the use of the price deflator does not make a significant difference in the results.

workers (individuals working for more than 35 hours in the reference week).[4] For further details on the characteristics of the surveys and the construction of the variables, see Appendix Appendix A.1.

Earnings inequality as summarized by the 90/10 log wage differential declined during the 2000s in Argentina and Chile,[5] reversing the increasing trend documented for the 1980s and the first years of the 1990s (Figure 1).[6] Earnings inequality peaked around 1996 in Chile and 2002 in Argentina. Inequality in Brazil declined steadily during the period of analysis, which starts in 1990. The contraction of the earnings distribution is quite significant: Since 2000, the ratio between the 90th and 10th percentiles contracted by 20 percent in Argentina, 28 percent in Chile, and a remarkable 46 percent in Brazil. Nonetheless, the levels of earnings inequality were among the highest in the world in 2013. The 90/10 log earnings ratio was close to 1.7 in the three countries, implying that the hourly earnings of a worker at the 90th percentile of the distribution is more than 5.5 times what a worker at 10th percentile earns. As a point of comparison, the OECD average of the 90/10 interdecile ratio was 4 in 2012.[7]

For the three countries, we observe a drop in the 90/50 and 50/10 log wage ratios after inequality peaked. This contrasts with a large body of literature from developed countries that shows that most of the changes in the wage structure have taken place at the top of the income distribution.[8] With the exception of Argentina, which displays a small increase in the 50/10 log wage ratio since the initial levels in 1995, the recent decline of all the inequality measures has brought inequality below the levels of the early 1990s.[9]

These changes in the distribution of earnings are taking place when

[4]We show in the robustness section of the paper that our main results are unchanged if we restrict the sample to prime-age workers (between ages 25 and 55) or include part-time workers. The average share of workers who reported working fewer than 35 hours per week is 15 percent in Chile and Brazil, and close to 28 percent in Argentina.

[5]See also Ferreira et al. (2008); Kahhat (2010); López-Calva and Lustig (2010); Gasparini and Lustig (2011); Gasparini et al. (2011a); Levy and Schady (2013); Lustig et al. (2013).

[6]See Cragg and Epelbaum (1996); Londoño and Szekely (2000); Sanchez-Paramo and Schady (2003); Behrman et al. (2007); Cornia (2010); Manacorda et al. (2010) among others.

[7]See `http://stats.oecd.org/`.

[8]See Katz and Autor (1999); Autor et al. (2005, 2008); Lemieux et al. (2009); Acemoglu and Autor (2011) and the references therein.

[9]Note that, unlike in Brazil and Chile, in Argentina there was a major economic crisis in 2001 that resulted in a rapid GDP contraction and a massive increase of unemployment, possibly affecting the evolution of inequality as highlighted by López-Calva and Lustig (2010). The study of these macroeconomic forces is beyond the scope of this paper, but we briefly return to this issue in Section 6.3.

the skill and demographic composition of the workforce is also changing significantly. Major shifts include changes in the education, age, and gender composition of the labor force. The share of workers with a primary education degree in the early 1990s was 45.7 percent in Argentina, 49.8 percent in Chile, and 76.4 percent in Brazil.[10] By 2013 that share had dropped in a range from 16.1 percentage points in Argentina to 31.5 percentage points in Brazil (Table 1). The gains in schooling are reflected in the increased shares of workers with high school education completed and those with a college degree. For example, the share of college-educated workers increased from 19.9 to 29.1 percent in Argentina, almost doubled in Chile (14.8 to 27.3 percent), and almost tripled in Brazil (7.8 to 19.5 percent).

Worker's average age increased by more than a year in Argentina (37.3 to 39), three years in Brazil (34.1 to 37.4), and four years in Chile (35.8 to 40.4). Even with the sharp rise in the levels of schooling, this demographic shift has resulted in a rise in the average level of potential experience.[11] This is especially significant in the case of Chile, where the share of workers with more than 20 years of potential experience increased by 11 percentage points.

In the early 1990s, female labor force participation was as low as 35 percent in Chile and close to 50 percent in Argentina and Brazil. By 2013 half the women between ages 16 and 65 in Chile were working, and female labor force participation in Argentina and Brazil was close to 60 percent. As a consequence of this shift, the employment share of women rose by more than 8 percentage points in Brazil and Chile.

3. Inequality, Workforce Composition, and Wage Structure

Changes in the composition of the labor force will affect inequality across and within labor market skill groups. Perhaps the most studied characteristic is education. Facilitating access to education to the poor is a powerful tool for social mobility and may lead to lower inequality in the long run. However, because the education premium is convex, in the short and medium term, educational upgrading may increase between-group inequality. This "paradox of progress" (Bourguignon et al., 2005) may occur even when changes in educational attainment are moderately in favor of low socioeconomic background groups, and was recently confirmed for several Latin American countries by

[10] See Appendix Appendix A for details on the aggregation of workers with incomplete levels of schooling.

[11] We define potential experience as: age-years of education-6.

Battistón et al. (2014). Moreover, within-group dispersion is typically much higher among highly educated workers, pushing inequality up when educational attainment increases (Lemieux, 2006).

The role of changes in the composition of the labor force associated with labor market experience and gender is less straightforward. Returns to experience are concave (Murphy and Welch, 1990), a force that would push between-group inequality down when the labor force is aging. However, within-group wage dispersion is higher among high-experience workers than among their low-experienced counterparts, possibly limiting the decline of inequality (Lemieux, 2006). Similarly, the importance of the composition effects associated with gender depends on the skills distribution of the women who are increasingly accessing the labor market.

A simple decomposition exercise can help disentangle the importance of composition and price effects for inequality. The exercise exogenously fixes the structure of relative wages at the average level across the past two decades and quantifies the counterfactual levels of the interquantile wage ratios under the observed compositional changes. Alternatively, we can keep the composition of the labor force fixed at a given point in time and construct counterfactual wage distributions to evaluate how changes in the schooling, experience, and male premiums have affected the observed inequality dynamics.

The decomposition we propose follows Firpo et al. (2007, 2009). They show that the properties of recentered influence functions (RIF) can be used to extend the traditional Oaxaca-Blinder decomposition to analyze distributional statistics beyond the mean (e.g., quantiles). As a starting point, consider a transformed wage-setting model of the form

$$RIFq_{\tau t} = X_t' \gamma_{\tau t} + \epsilon_{\tau t} \;\; \text{for} \;\; t = 1, 0, \tag{3.1}$$

where t identifies the initial ($t = 0$) and final ($t = 1$) periods; $RIFq_{\tau t}$ represents the value of the RIF corresponding to the τth quantile of the earnings distribution at time t; $\epsilon_{\tau t}$ is the projection error; and X_t is a vector of socio-demographic characteristics including quadratic terms in education and experience and a female dummy.[12] We can estimate equation (3.1) by ordinary least squares (OLS) and express the estimated difference over time of the expected value of the wage quantile $\hat{q}_\tau$ as

$$\Delta\hat{q}_\tau = \underbrace{\left(\overline{X'}_1 - \overline{X'}_0\right)\hat{\gamma}_P}_{\Delta\hat{q}_{X,\tau}} + \underbrace{\overline{X'}_{\tau P}\left(\hat{\gamma}_{\tau 1} - \hat{\gamma}_{\tau 0}\right)}_{\Delta\hat{q}_{S,\tau}}, \tag{3.2}$$

where overbars denote averages, and $\hat{\gamma_{\tau P}}$ and $\overline{X}_P$ correspond to the estimated vectors of parameters and the explanatory variables of a wage-setting model in which observations are pooled across the periods.[13] Here, $\hat{q}_{X,\tau}$ corresponds to the composition effect, which captures the part of the change in the τth wage quantile that is accounted for by changes in the average skill-demographic characteristics of the workforce, given that we set the skill returns at their (weighted) average over the two periods; and $\hat{q}_{S,\tau}$ is the wage structure effect, capturing how changes in returns affect wages at the quantile τ, given that the observable characteristics are fixed to be equal to their (weighted) average over time. The model is additive. Hence, differences over time between quantiles a and b can be simply expressed as:

$$\underbrace{\Delta\hat{q}_a - \Delta\hat{q}_b}_{\text{Overall}} = \underbrace{(\Delta\hat{q}_{X,a} - \Delta\hat{q}_{X,b})}_{\text{Composition}} + \underbrace{(\Delta\hat{q}_{S,a} - \Delta\hat{q}_{S,b})}_{\text{Wage Structure}} \tag{3.3}$$

We focus on the determinants of changes in three inequality measures: the interquantile log earnings ratios 90/10, 90/50, and 50/10. The results are shown in Table 2. For the three countries, we observe a very similar pattern: changes in the skill and demographic composition of the workforce have had an unequalizing effect on the distribution of earnings as measured by the log 90/10 wage ratio. In Argentina, the counterfactual change is relatively small (5.7 percent increase), but it is sizable in Chile (28.3 percent increase) and Brazil

[12] As in the Oaxaca-Blinder decomposition for the mean, this decomposition is not invariant to the reference variable chosen when the covariates are categorical. We limit this problem by using quadratic polynomials in years of education and potential experience. In the case of gender, we repeated the decomposition with a male dummy and a female dummy. The results were qualitatively similar.

[13] This specific counterfactual allows us to analyze composition and wage structure effects relative to a baseline defined by the (weighted) mean returns and (weighted) mean characteristics over the two periods, eliminating the interaction term present in other decompositions (Oaxaca and Ransom, 1994).

(32.6 percent increase).[14] These unequalizing effects of compositional changes are observed at both ends of the earnings distribution, but the magnitude tends to be larger in the upper half (90/50 log wage ratio). In Brazil and Chile, compositional changes would have pushed up the 90/50 wage ratio by 25 and 20 percent, respectively. Thus, changes in the skill-demographic composition contributed to the increase in inequality in the 1990s in Argentina and Chile, but cannot account for the subsequent decline in these two countries or the reduced inequality in Brazil.

Wage structure effects are key to understanding the evolution of inequality. Earnings inequality would have declined by 13 percent had changes in the composition of the labor force been kept constant in Argentina, and as much as 41 percent in Chile and 67 percent in Brazil. Of course, these are partial equilibrium counterfactuals, which do not take into account the impacts compositional changes may have had on the returns to observable characteristics, an aspect we will address below.

Among the wage structure effects, changes in the schooling and experience premiums had a prominent role in the observed inequality trends in the three countries. Changes in the schooling premium account for a decline of the interquantile 90/10 log ratio of 0.271 log points in Argentina, 1.153 in Brazil, and 1.685 in Chile. The wage structure effects associated with education outweighed the compositional effects. Overall, the total contribution of education to the decline of the 90/10 log ratio was -.216, -.901, and -1.474 log points in Argentina, Brazil, and Chile, respectively. Compared with the contribution of education, the contribution of changes in the experience premium to the decline of the 90/10 log ratio is of similar magnitude in Argentina and Brazil (-.282 and -.825), and somewhat smaller in Chile (-.497). The importance of changes in the experience premium for the decline in inequality contrasts with the focus in the academic literature on the evolution of the schooling premium. Finally, although changes in the gender wage gap also had equalizing effects, their impact on overall inequality trends was much smaller than that of experience and education.

The contribution of education to the decline in inequality is entirely driven by changes in the lower tail of the distribution in Argentina and Brazil, while its impact on the 90/50 gap is not statistically significant. This is particularly strong in Brazil, where the contribution of the schooling premium to

[14] All percent changes are calculated by taking the exponential of the respective values, which are expressed in logarithms, and subtracting one.

the decline of the 50/10 gap (-1.228 log points) is by far the most significant factor. As we shall see below, this is also consistent with the much sharper decline of the high school premium than the college premium in Brazil. In Chile education contributed to a compression of wages at the two ends of the distribution. The contribution of the experience premium to the decline of inequality is visible in the two halves of the distribution, except for Argentina, where it only contributed to a decline of the 90/50 gap (-.253 log points).

The decomposition exercises show that the observed patterns in earnings inequality are mostly driven by how the wage structure changed over time; the exercises give no indication as to why those relative returns changed. Moreover, the wage structure effects are calculated under a counterfactual in which the skill-demographic composition of the workforce is held constant, which we know was not the case. A natural hypothesis, then, is that the wage structure is changing because of the compositional changes, not in spite of them. This would be the case if workers with different skill-demographic characteristics are not perfectly substitutable in production, so that changes in relative supplies directly influence relative wages. In the next section, we provide descriptive evidence that this simple mechanism is consistent with the observed trends. We then proceed to formulate a model that can rationalize the patterns in the data.

4. Skill Supply and Demand and the Evolution of Relative Returns

The previous section underscored the importance of changes in the education and experience premiums for changes in inequality. Falling schooling and experience premiums appear as primary factors in the recent evolution of wage dispersion. We start this section by laying out basic facts about the evolution of these two premiums in each country. To isolate changes in the premiums, we present composition-adjusted series of relative wages (see Figure 2). In particular, we first compute mean (predicted) log real earnings in each country-year for 70 skill-demographic groups (five education levels, seven potential experience categories in five-year intervals, males and females). The mean wages for the broader groups shown in the figures are then calculated as fixed-weighted averages of the relevant subgroup means, where the weights are equal to the mean employment share of each subgroup across all years. This adjustment ensures that the estimated premiums are not mechanically affected by compositional shifts.

The first column in Figure 2 shows the evolution of the college/high school and high school/primary log wage gaps for full-time workers. The high school/primary wage gap follows quite closely the evolution of earnings inequality. The peaks in overall (90/10) inequality coincide with the peaks in the high school/primary wage gap in Argentina and Chile. In Brazil, both series have fallen since the start of the sample. The decline in the high school/primary wage gaps is substantial. After the peak, the high school premium declined by 12 percent in Argentina, 19 percent in Chile, and 46 percent in Brazil. The college/high school wage gaps also fell during the same period, although at a slower pace in Brazil and Chile. In Argentina, where the expansion of secondary education took place earlier and employment of college graduates quickly gained ground with respect to high school graduates (see Table 1), the reduction of the college premium (22 percent since 2002) is even larger than the decrease in the high school/primary wage gap. This is consistent with the fact documented above that the reduction of inequality in Argentina was stronger in the upper tail (log 90/50 earnings gap) than in the lower half of the distribution (log 50/10 earnings gap).

Column 2 in Figure 2 shows the evolution of the composition-adjusted wage gap between high-experience (20-29 years of potential experience) and low-experience (0-9) workers across three educational groups: college, high school, and primary completed. The experience wage gap (20-29/0-9 log wage ratio) declined in the three countries and for all education groups, but this was not a unitary phenomenon. The reduction of the experience premium was concentrated among workers with different education levels, depending on the country.

In Argentina, the (20-29/0-9) log wage ratio falls for the three groups, but much more markedly among college-educated workers. The (20-29/0-9) log wage ratio among college educated workers declined by 26 log points, after peaking in 1997, compared with a reduction of the same ratio of 16 log points among high school graduates and stagnation on the ratio for workers with primary education. Real wage growth throughout the period was negative among college educated workers with more than 10 years of experience, but it was positive for the least experienced (see Table Appendix A.1). In sum, the decline of relative wages among college educated workers that was documented for Argentina has a strong experience gradient: it was concentrated among more experienced workers.

In Brazil and Chile, the fall of the experience premium is stronger among workers with completed high school education. Throughout the period, the

(20-29/0-9) log wage ratio among high school educated workers declined by 47 log points in Brazil and 35 log points in Chile. The experience premium of workers with primary education or college also fell, but at a lower rate. In Brazil (Chile), the (20-29/0-9) wage gap fell by 23 (19) log points among workers with at most primary completed, and by 34 (22) log points among college graduates.

Our main hypothesis is that changes in the schooling and experience premiums over the past two decades have responded to movements in the relative supply of workers. Table 3 contains suggestive evidence. For example, for high school-educated workers in Chile, although wages for all workers grew, there is an inverse relationship with worker experience. The wages of the least experienced (0 to 9 years of experience) workers increased by 58 log points over the period, while those of workers with more than 29 years of experience increased by just 8 log points. Changes in log working-age population go in the opposite direction. Workers with a high school degree increased overall, but the expansion is much stronger among workers with high experience: 145 log points among workers with more than 29 years of experience against 44 log points for workers with 0 to 9 years of experience. Similar patterns are found in Argentina and Brazil and among workers with other levels of schooling.

As a further motivation for the empirical analysis that we develop in the next section, Figure 3 shows the comovement between changes in log wages and log labor supply across education/experience groups.[15] In each case, we have de-trended the series of log changes from a country-specific linear trend, to capture possible secular movements such as skill biased technical change or demand shifts favoring a particular skill group. Panel (a) shows a very tight relationship between the high school/primary wage gap and the corresponding log change in labor supply. The slope coefficient is $-.53$ and is highly significant (s.e. 0.05). The relationship between labor supply and relative wages is negative but weaker when we consider changes in the log college/high school wage gap and a corresponding measure of labor supply (panel (b)), but the slope of a linear fit of the cross-plot remains statistically significant (coeff: -0.24; s.e. 0.08).

Panels (c) and (d) in Figure 3 show the relationship between de-trended experience premiums and their corresponding relative labor supply measures

[15] Following the analysis in Section 5, we use a broad measure of labor supply, the working-age population of each education/experience group, but the results are very similar using log employment changes.

among unskilled workers (those with primary and secondary education, in panel (c)) and college educated workers (panel (d)).[16] In the two cases, the cross-plots indicate a clear negative relationship, confirming the previous discussion and the potential importance of labor supply changes for the evolution of the experience premium. The relationship is stronger among unskilled workers than among college educated workers, but in both graphs the linear fit of the cross-plot is negative and statistically significant.[17]

Relative quantities and relative prices appear to move in opposite directions for the two main drivers of changes in the wage structure: education and experience. To what extent can movements in the schooling and experience premiums be explained by changes in labor supply? The answer to this question depends on the degree of substitutability across skills. Thus, we move now to delineate a stylized model of the supply and demand of skills that allows us to estimate the impact of changes in labor supply on the skills gap.

5. Supply-and-Demand Model

The basic framework follows the canonical work of Katz and Murphy (1992), Murphy and Welch (1992), and Katz and Autor (1999). Workers are divided into skill groups, which are allowed to be imperfect substitutes in production. In particular, we assume that aggregate production in this economy can be described by a multilevel nested CES function. At the top level, output is produced as a CES combination of labor with high (college education completed or more) and low (high school degree at most) skills,

$$Y_t = \lambda_t \left(L_{Ut}^{\rho} + \alpha_t L_{Ct}^{\rho} \right)^{1/\rho}, \tag{5.1}$$

where Y_t is total output at time t; L_{Ut} is the total supply of low-skill labor; L_{Ct} is the total supply of college-educated workers; λ_t is a scale parameter that is allowed to vary over time to capture skill-neutral technological change; α_t

[16] We follow the analysis laid out in Section 5 and consider changes in log relative wages (and labor supply) for: workers with $10-19$ years of experience versus workers with $0-9$ years of experience; workers with $20-29$ years of experience versus workers with $0-9$ years of experience; and workers with 29 years of experience or more versus workers with $0-9$ years of experience. Panel (c) in Figure 3 shows each log wage and labor supply change measure for workers with primary and secondary education separately, while panel (d) focuses on college-educated workers.

[17] Among high school and primary educated workers, the slope of the linear fit is -0.35 (s.e. 0.027), and for college educated workers the slope is -0.16 (s.e. 0.05).

is a time-varying parameter that captures differences in relative productivities and relative demand between college and unskilled labor; and ρ is a function of the elasticity of substitution (σ_ρ) between college-educated and unskilled labor: $\sigma_\rho = \frac{1}{1-\rho}$.

As noted by Katz and Autor (1999), that we model the economy using an aggregate production function means that we have to be careful not to interpret the parameters as if we were dealing with individual firms. For example, the elasticity of substitution σ_ρ reflects not only technical substitution possibilities between workers at the firm level, but also outsourcing and substitution across goods and services in consumption. In a similar way, α_t captures relative productivity changes at the intensive (workers performing better at their current jobs) and extensive margins (e.g., a shift in work tasks across workers of different skill groups), changes in relative prices or quantities of non-labor inputs, and shifts in product demands among industries with different skill intensities.

Following Manacorda et al. (2010), we further divide the total supply of unskilled labor (L_{Ut}) into two subgroups. The first subgroup is formed by labor from workers who have at least obtained a high school degree, but who have not completed any post-secondary education (L_{Ht}). The second subgroup comprises labor from workers who have at most obtained a primary education degree (L_{Pt}).[18] The aggregation is done using a productivity-weighted CES combination of the form

$$L_{Ut} = \left(L_{Pt}^{\delta} + \beta_t L_{Ht}^{\delta}\right)^{1/\delta}, \tag{5.2}$$

where β_t is a time-varying parameter that captures differences in relative productivities between the two subgroups and changes in relative demands; and δ is a function of the elasticity of substitution (σ_δ) between the two low-skill types.

The descriptive evidence showed that changes in the experience premium were important in the three countries, and not necessarily parallel across schooling levels. Hence, we follow Card and Lemieux (2001) and Manacorda et al. (2010) and relax the assumption of perfect substitution across experience groups by assuming that the supply of college, high school, and primary educated workers depends on a CES sub-aggregate across experience levels:

[18] Hence, high school dropouts are included in this group.

$$L_{Kt} = \sum_i \left(\phi_{Kit} L_{Kit}^{\theta_K} \right)^{1/\theta_K} \quad \text{for} \quad K = P, H, C \ , \tag{5.3}$$

where θ_K are functions of the partial elasticities of substitution σ_{θ_K} across experience subgroups within skill level K, and ϕ_{Kit} are relative efficiency parameters for each potential experience group within skill level K. In contrast with Card and Lemieux (2001) and Manacorda et al. (2010), these relative productivities are time-varying. This allows for forces other than changes in relative labor supply to determine the evolution of the experience premium.

5.1. Empirical Implementation

We assume that the economy is operating along the competitive equilibrium demand curve. The implication is that the wage of each labor type is fully determined by its marginal productivity. Further, we assume that the labor supply of each labor type is exogenously determined. This is a strong assumption, especially since we observe a sharp movement of women into the labor market during the past 25 years. To mitigate problems of selection arising from endogenous labor supply responses that differ across skill groups, the empirical implementation of the model uses working-age population as a labor supply measure, irrespective of employment status or hours worked. By contrast, relative wages across skill groups are measured for full-time male workers ages 16-65, which eliminates problems of selection among female workers.[19]

Our baseline specification divides potential experience into four groups. We denote group 1 as the least experienced workers ($0-9$ years of experience), group 2 as workers with moderate ($10-19$) years of experience, group 3 as experienced workers ($20-29$), and group 4 as the most experienced (more than 29 years of experience). Thus, we have 12 different labor types in the model (3 schooling levels $\times$ 4 potential experience groups). To reduce the parameter space, we assume that the elasticity of substitution and relative productivity parameters within the unskilled group (workers with primary and high school education) are the same, but are allowed to differ from the respective parameters for college educated workers.

In this setting, the marginal product of labor of each experience/education group depends on its own group labor supply and the aggregate supply of la-

[19] In the robustness section of the paper, we show that the results are robust to alternative definitions of labor supply and relative wages.

bor for its level of education. Denoting lowercase variables as the natural logarithms of the respective uppercase variables, the eight equilibrium conditions for unskilled workers are summarized in the following expression:

$$w_{Kit} = \lambda_t + \log \zeta_{Kit} + \frac{1}{\sigma_\rho}(y_t - l_{Ut}) + \frac{1}{\sigma_\delta}(l_{Ut} - l_{Kt}) + \frac{1}{\sigma_{\theta_U}}(l_{Kt} - l_{Kit}),$$
$$\text{for} \quad K = H, P \quad \text{and} \quad i = 1, 2, 3, 4 \quad (5.4)$$

where $\zeta_{P1t} = 1$; $\zeta_{P2t} = \phi_{U2t}$; $\zeta_{P3t} = \phi_{U3t}$; $\zeta_{P4t} = \phi_{U4t}$; $\zeta_{H1t} = \beta_t$; $\zeta_{H2t} = \beta_t \phi_{U2t}$; $\zeta_{H3t} = \beta_t \phi_{U3t}$; and $\zeta_{H4t} = \beta_t \phi_{U4t}$. In a similar way, the equilibrium conditions for the four experience groups among college educated workers are summarized by:

$$w_{Cit} = \lambda_t + \log \zeta_{Cit} + \frac{1}{\sigma_\rho}(y_t - l_{Ct}) + \frac{1}{\sigma_{\theta_S}}(l_{Ct} - l_{Cit}), \text{ for } \ i = 1, 2, 3, 4 \quad (5.5)$$

where $\zeta_{C1t} = \alpha_t$; $\zeta_{C2t} = \alpha_t \phi_{C2t}$; $\zeta_{C3t} = \alpha_t \phi_{C3t}$; and $\zeta_{C4t} = \alpha_t \phi_{C4t}$.

The model has two types of relevant parameters that we wish to estimate: four parameters that are functions of the elasticities of substitution between labor groups (ρ, δ, θ_U, and θ_C), and after normalizing $\phi_{U1t} = \phi_{C1t} = 1$, a set of time-varying relative efficiency/demand shifter parameters (α_t, β_t, ϕ_{U2t}, ϕ_{U3t}, ϕ_{U4t}, ϕ_{C2t}, ϕ_{C3t}, and ϕ_{C4t}). As shown by Johnson and Keane (2013), we could fit the trends in relative wages perfectly if we did not impose any restrictions on the evolution of the relative efficiency parameters, but this would mean that we would not be able to identify the elasticities of substitution. We then restrict in the baseline specification these relative productivities to follow a cubic trend in their natural logarithm, and assess the sensitivity of the results to this assumption in the robustness section.[20]

The parameters of the model are estimated in three sequential steps. In the first step, we derive from the equilibrium condition (5.4) a set of expressions that relate the evolution of relative earnings across experience groups to their relative supplies. In particular, we have eight equilibrium conditions of the general form:

[20] We also tried quartic time trends, without significant changes to our main results. The estimated parameters associated with the fourth order of the quartic specification were no longer statistically significant.

$$w_{Kit} - w_{Kjt} = \phi_{Uit} - \phi_{Ujt} - \frac{1}{\sigma_{\theta_U}}(l_{Kit} - l_{Kjt}),$$
$$\text{for} \quad K = P, H \;\; ; \;\; i, j = 1, 2, 3, 4 \quad \text{and} \quad i > j. \quad (5.6)$$

Similarly, from equation (5.5) we obtain four additional equilibrium conditions for college-educated workers:

$$w_{Cit} - w_{Cjt} = \phi_{Cit} - \phi_{Cjt} - \frac{1}{\sigma_{\theta_C}}(l_{Cit} - l_{Cjt}),$$
$$\text{for} \quad i, j = 1, 2, 3, 4 \quad \text{and} \quad i > j. \quad (5.7)$$

Using equations 5.6 and 5.7, we estimate the elasticities of substitution across experience groups for unskilled (σ_{θ_U}) and college-educated workers (σ_{θ_C}), and the corresponding demand shifters.[21] These estimates are used to construct the aggregate labor supply measures of primary (P_t), high school (H_t), and college (C_t) educated workers.

In the second step, we estimate the parameters governing the elasticity of substitution between the two unskilled labor groups (σ_δ), and the relative demand for high school educated workers (β_t). This is obtained from a set of relative wage equations derived from conditions (5.4) that follow the general form:

$$w_{Hit} - w_{Pit} = \beta_t - \frac{1}{\sigma_\delta}(l_{Ht} - l_{Pt}) - \frac{1}{\sigma_{\theta_U}}\left[(l_{Hit} - l_{Ht}) - (l_{Pit} - l_{Pt})\right]$$
$$\text{for} \quad i = 1, 2, 3, 4. \quad (5.8)$$

The first term in these equations relates the high school premium to aggregate relative labor supplies, and provides an estimate of σ_δ. But the high school premium is allowed to differ across workers with different experience. The second term relates the high school premium of workers with experience level i to the deviations of the supply of that specific experience group with respect to the aggregate supply of workers with primary and secondary education. This provides a second estimate of the elasticity of substitution between

[21] Normalizing $\phi_{U1t} = \phi_{C1t} = 1$, we estimate ϕ_{U2t}, ϕ_{U3t}, ϕ_{U4t}, ϕ_{C2t}, ϕ_{C3t}, and ϕ_{C4t}.

experience subgroups among unskilled workers, serving as an internal consistency check. In principle these estimates should be similar to those obtained in the first step.

Using the estimated elasticities and relative productivity parameters from the first two steps, we construct a series of supply of unskilled labor (U_t). The third estimation step derives from equations (5.4) and (5.5) the following expressions:

$$\log\left(\frac{\widetilde{W_{Cit}}}{W_{Kit}}\right) = \log\alpha_t - \frac{1}{\sigma_\rho}\log\left(\frac{L_{Ct}}{L_{Ut}}\right) - \frac{1}{\sigma_\delta}\log\left(\frac{L_{Ut}}{L_{Kt}}\right) - \frac{1}{\sigma_{\theta_C}}\log\left(\frac{L_{Cit}}{L_{Ct}}\right)\dots$$
$$-\frac{1}{\sigma_{\theta_U}}\log\left(\frac{L_{Kt}}{L_{Kit}}\right) \quad \text{for} \quad K = H, P \quad \text{and} \quad i = 1,2,3,4, \quad (5.9)$$

where the terms $\log\left(\frac{\widetilde{W_{Cit}}}{W_{KJt}}\right)$ are the wage gaps of college/unskilled workers of experience group i that have been "demand-detrended" using the efficiency parameters estimated in the previous steps.[22]

The third step estimates all the parameters of the production function. Hence, the estimation of equations 5.9 provides a third estimate of σ_{θ_U}; second estimates of σ_{θ_C} and σ_δ; and a first estimate of the elasticity of substitution between skilled and unskilled workers σ_ρ.

The model is estimated with the pooled data and each estimation step includes country fixed effects and equation fixed effects. In the baseline model, we allow the relative efficiency parameters to be country-specific but restrict the parameters that govern the elasticities of substitution to be common across countries. Because the second and third steps invoke generated regressors, the standard errors are bootstrapped. This procedure contrasts with Card and Lemieux (2001), Manacorda et al. (2010), and Autor et al. (2008), who report conventional standard errors. As it turns out, the distinction is not trivial. When we compare our estimates of the standard errors with the conventional robust ones, we often find that the bootstrapped errors are up to two times larger. Hence, it seems that taking into account generated regressors is crucial for correct inference in this context.

[22]In particular, $\log\left(\frac{\widetilde{W_{Cit}}}{W_{Hit}}\right) = \log\left(\frac{W_{Cit}}{W_{Hit}}\right) - \log\frac{\hat{\phi}_{Cit}}{\hat{\beta}_t\hat{\phi}_{Uit}}; \log\left(\frac{\widetilde{W_{Cit}}}{W_{Pit}}\right) = \log\left(\frac{W_{Cit}}{W_{Pit}}\right) - \log\frac{\hat{\phi}_{Cit}}{\hat{\phi}_{Uit}}$ for $i = 1,2,3,4$, $\hat{\phi}_{C1t} = 1$ and $\hat{\phi}_{U1t} = 1$.

6. Results: Labor Supply, Experience and Education Premiums

6.1. Labor Supply and the Experience Premium

The results of the estimates in the first step for unskilled workers are shown in column 1 in Table 4. The fit of the model is good, with $R^2 = 0.9$. The coefficient of the relative supply of experience (-0.28, with s.e. 0.017) confirms the importance of changes in the relative supply of experience for movements in the experience premium. The implied elasticity of substitution among experience groups within unskilled workers (σ_{θ_U}) is 3.6 and highly significant. The second and third steps provide additional estimates of the elasticity of substitution across experience groups, ranging from 3.1 (column 3) to 4.1 (column 4). All the estimates are statistically significant and equality of the coefficients across the different specifications cannot be rejected, which provides an internal consistency check. These estimates are somewhat lower than those reported in Card and Lemieux (2001), who find an elasticity of substitution across ages in Canada, the United Kindom, and the United States in the range of 4 to 6. Manacorda et al. (2010) cannot reject perfect substitution across age groups in a sample of Latin American countries during the 1980s and 1990s. However, Card and Lemieux (2001) and Manacorda et al. (2010) impose a common elasticity of substitution for low- and high-skilled workers and constant relative productivity parameters across age groups, and hence are not strictly comparable.

The first column in Figure 4 shows the evolution of the unskilled experience premium for workers with high $(20-29)$ experience: the $20-29/0-9$ experience wage gap. It also shows two predictions from the model. The first prediction sets the efficiency parameters to 1 and illustrates how far the model goes in explaining the experience premium for this particular group invoking supply changes alone. The second prediction uses the full empirical model, and serves as a visual test of the goodness of fit. Based on the observed changes in relative supplies in each country and $\hat{\sigma}_{\theta_U}$, the predicted fall in the experience premium absent any demand changes is 8.3 log points in Argentina, 11.3 log points in Brazil, and 14.7 log points in Chile. Compared with observed reductions of 16.9 , 26.6, and 28.4 log points in the three countries, this indicates that rising experience in the labor market explains between 42 and 52 percent of the fall in the $20-29/0-9$ experience wage gap.

The first column in Figure 5 shows the estimated relative efficiency trends across experience groups (ϕ_{Ujt}). The general pattern is similar for the premiums of moderate (10-19), high (20-29), and very high (>29) experienced workers in the three countries. The demand shifters are fairly flat, showing that relative supply by itself does a decent job in explaining the movements of relative wages among experience groups. However, the shifters also show a mild downward trend during the 2000s, which indicates that during the past period demand forces (or technical change) have moved against experienced workers among the unskilled.

Column 2 in Table 4 shows the results of estimating the elasticity of substitution across experience groups for college-educated workers. The coefficient of relative labor supply of experience is $-.093$, but the precision of the estimate is low and we cannot reject the null hypothesis that experienced and inexperienced workers among college graduates are perfect substitutes. When we run the third step, in column 4, the coefficient of relative supply of experience for college graduates doubles ($-.181$) and becomes statistically significant, with an implied elasticity of 5.5.

As it turns out, the lack of precision in the estimation of the elasticity of substitution among college graduates in the first step is to a great extent driven by the behavior of very experienced workers in Brazil. Among Brazilian male college graduates, the experience premium with respect to workers with $0-9$ years of experience declines by 4.6 log points for workers with $10-19$ years of experience and by 26.6 log points for workers with $20-29$ years of experience. This pattern is similar to what we observe in Argentina and Chile (Table Appendix A.1). However, the experience premium of workers with more than 30 years of experience increases by 6.1 log points. This contrasts with Argentina and Chile, where the experience premium monotonically deteriorates as experience increases. When we dummy out this group from the analysis in the first step, the estimate of relative labor supply of experience becomes $-.153$, much closer to the estimate in the third step, and is statistically significant at the 5 percent level (s.e. 0.073).

This exceptional behavior of very experienced workers in Brazil may be related to selection effects associated with early retirement. The Brazilian pension system creates strong incentives for early retirement, especially among highly educated workers who have held formal jobs for most of their careers (Leme and Málaga, 2001). These incentives increased over the period of study (Queiroz, 2018). While the hazard rate of retirement in OECD countries for workers ages 50-60 is virtually zero, it ranges from 5 to 10 percent in the case of Brazil. Moreover, labor force participation rate among males ages 55 and older declined over the past decades, and presents a negative relationship with years of schooling (Queiroz, 2018). In case of positive selection, the experience premium of college graduates with more than 30 years of experience may be overstated.

The impact of relative labor supply trends on the evolution of the experience premium of college-educated workers is weaker than the impact observed among unskilled workers. Using the estimates in column 2 in Table 4, the effect is virtually 0, as evidenced from the predictions in Figure 4. But even if we use the estimates from the third step (column 4 in Table 4), the impact is fairly limited. Table Appendix A.1 shows that the log $20-29/0-9$ experience wage gap of college-educated workers declined 23.4, 26.6, and 15.6 log points in Argentina, Brazil, and Chile, respectively. The predicted relative labor supply effects corresponding to the estimates in column 4 are reductions of 4.8, 6.9, and 7.2 log points in each country. Hence, relative labor supply explains at most 20 to 33 percent of the observed decline of the $20-29/0-9$ experience wage gap.

6.2. Labor Supply and the Schooling Premium

The estimates from the second step assess the effect of the relative supply of workers with secondary education with respect to primary on the high school premium. The estimated coefficient of relative supply is -0.43 (column 3 in Table 4). This is virtually identical to the coefficient obtained in the third step (column 4). Both estimates are statistically significant and imply an elasticity of substitution across these two groups (σ_δ) of 2.3. This estimate is also very close to the elasticity of substitution across these two groups reported by Manacorda et al. (2010) for $1980-1990$, which is remarkable, considering differences in the models, periods of analysis, and samples of countries.

Changes in the relative supply of labor do a fairly good job in fitting the observed dynamics of the high school premium, especially in Brazil where the fit is almost perfect. Labor supply changes predict a massive decline of the high school premium of -68.4 log points in Brazil, while the observed decline was -71.9. The prediction of wage changes associated with movements in relative supply trace the observed evolution of the high school premium very closely (see Figure 6). This is confirmed by the estimates of the residual demand shifter, which are very close to zero throughout the period (column 1 in Figure 7). Thus, the prominent role of education in the decline of lower tail inequality (log 50/10 percentile ratio) that was documented earlier was fundamentally driven by the massive increase of high school-educated workers with respect to those who just completed primary education: an expansion of

65.6 log points in Argentina, 85.3 log points in Chile, and 135.3 log points in Brazil.

In Argentina and Chile, the increase in the relative supply of high school-educated workers actually overshoots the observed decline in the high school premium. The decline of the premium is much smaller than in Brazil (-12.7 and -31.1 log points in Argentina and Chile, respectively). Even if the expansion of the relative supply of high school-educated workers is not as large as in Brazil, the reduction of the high school/primary log wage gap predicted by changes in supply was -31.9 and -51.1 log points in the two countries. Investigating the dynamics reveals an important difference between the two countries. In Chile, the residual demand shifter swiftly slopes upward during the first half of the 1990s and flattens out during the rest of the period (Figure 7). Thus, it was fundamentally a sharp increase in the demand for high school-educated workers during the early 1990s that prevented the high school premium from falling even more. In contrast to this episodic event, Argentina presents a mild upward trend throughout the period.

The final step provides an estimate of the impact of the relative supply of college-educated with respect to unskilled workers on the college premium (column 4 in Table 4). The estimated effect is -0.80 and highly significant, implying an elasticity of substitution (σ_ρ) of 1.25. The estimated elasticity is somewhat lower than previous estimates. Katz and Murphy (1992) and Johnson and Keane (2013) report an elasticity of substitution between college- and high school–educated workers of 1.5 for the United States, similar to the 1.6 reported by Goldin and Katz (2009). Card and Lemieux (2001) find estimates in the range of 2 to 2.5 in Canada and the United States; and Manacorda et al. (2010) in the range 2.6 to 5, depending on the specification.

As was the case with the high school premium, the changes predicted by labor supply of the college premium match the data quite well in Brazil, but miss important dynamics in Argentina and Chile. In Brazil, the college premium with respect to workers with primary and secondary education (unskilled) remained fairly stable during the 1990s, and started trending downward during the 2000s (Table 6). Overall, the college/unskilled wage gap declined 56.6 log points, while labor supply changes predicted a reduction of 67.1 log points.

In Argentina and Chile, changes in labor supply predict a sharper decline of the schooling premium than the one observed in the data (Figure 6). This is fundamentally due to a mismatch between a rapid increase in the college premium during the 1990s and a secular rise in the relative supply of college-educated workers (which translates into a predicted reduction of the premium) in the past two decades. Thus, the relative demand trend estimated in the model has a strong positive slope during the 1990s for the two countries (Figure 7). During the 2000s, the fit of the model with labor supply changes alone is better, as evidenced by a mild downward-sloping demand trend. From 2000 to 2013, the cumulative decline of the premium is 46.8 (32.8) log points in Argentina (Chile), and that predicted by changes in labor supply is 35.1(23.1) log points. Thus, it is fundamentally the strong increase in the demand for high-skill labor during the 1990s that sustained the college premium during this period. Labor supply changes would have predicted a strong decline. Those supply changes were sustained during the 2000s, but relative demand shifted downward, contributing to the decline of the college/unskilled wage gap.

6.3. Robustness

In this section, we present several exercises aimed at understanding how sensitive our results are to the modeling assumptions and sample decisions. We focus our discussion on the estimates from the third step, which include all the parameters of the relative demand shifts, but the results are consistent across specifications. The results are presented in Table 5, which includes in column 1 the results from our baseline specification to facilitate the comparison.

One criticism of the canonical supply-and-demand framework is the potential sensitivity of the elasticity estimates to how one approximates the unobserved changes in relative demands (Borjas et al., 2012). As a robustness check, we modify the baseline model so that instead of capturing the efficiency parameters with a third-order polynomial, we use year fixed effects that are common across the three countries. The results are shown in column 2 in Table 5. Although the point estimates are somewhat lower, the implied elasticities are very similar. In the working paper version of this paper (Fernandez and Messina, 2017), we experiment with a simplified version of the model with two experience groups: high (more than 20 years of experience) and low (less than 20 years of experience). The results are very similar to those reported here, except for the elasticity of substitution between college-educated and unskilled workers, which is slightly larger at 2.09.

Our baseline specification uses a broad measure of labor supply: log working-age population. The advantage of this measure is that it is likely to be predetermined, because changes in the education of the working-age population are the result of past human capital investments. One drawback is that relative wages across groups are likely to be less sensitive to inactive population, and the share of inactives may differ across skill groups. Column 3 presents the results when labor supply is approximated by the employed population in each cell, which does not suffer from this limitation. There are virtually no differences with respect to the baseline parameters, except for a small decline in the impact of the relative supply of skilled/unskilled workers. Column 4 presents the estimates when labor supply is calculated by adding up the total number of hours worked by each labor type. This is done to account for potential changes in the intensive margin. Since we only have information on hours worked for individuals who are employed, we impute the numbers for individuals outside the workforce by assigning them the average number of hours worked by an employed worker with the same education, potential experience, and sex in the respective country-year. The estimated impacts of the relative supply measures are very similar to the baseline estimates.

Column 5 includes only workers ages 25-55. The sharp changes in the educational composition of the workforce might lead to sample selection issues associated with a larger share of younger workers remaining in the education system, which would affect our wage series. Limiting the sample to prime age workers increases the impact of the relative supply of experience among unskilled workers, from -.245 to -.373, while the rest of the parameters are qualitatively similar to the baseline estimates. Column 6 shows that the estimates are very similar when we include part-time and full-time workers in the calculation of the wage series.

The supply-and-demand framework presented above is silent about the role of institutional and cyclical conditions in the labor market in explaining changes in the wage structure. The period of analysis includes sharp changes in cyclical conditions, with unemployment rising in the three countries during the late 1990s and early 2000s and then declining steadily after 2002. This is particularly marked in Argentina, which suffered a major economic and financial crisis between the end of 2001 and 2003, with the unemployment rate peaking above 20 percent. The decline of unemployment coincided with the so-called commodity super cycle, which brought about sharp terms-of-trade improvements in the three countries. Finally, changes in institutions were also important during this period. Between 1990 and 2012, the real hourly minimum wage increased by 120 percent in Brazil and by 138 percent in Chile. In Argentina it only increased after the 2001 crisis but at a fast pace, more than doubling in the decade that followed. We study the role of the minimum wage, unemployment, and terms of trade within our framework in the working paper version of this paper (Fernandez and Messina, 2017). We find that, depending on the country, these factors are important for the evolution of wage inequality, confirming results from other studies (Gasparini et al., 2011b; Cornia, 2012; Székely and Mendoza, 2015). However, the estimated elasticities of substitution are very similar to those reported in the baseline specification. The role of changes in labor supply for the evolution of the schooling and experience premiums remains unaltered.

7. Conclusions

After a decade of stagnant or rising earnings inequality, the distance between the top and bottom earners in Latin America fell sharply during the late 1990s and 2000s. This trend was in sharp contrast to the experience of developed countries during the same period. This paper has offered a detailed accounting of the main trends in three of the largest countries in the region: Argentina, Brazil, and Chile, and focused on two key drivers: the evolution of the schooling and experience premiums and their relationship with labor supply trends.

We started by building counterfactual distributions, keeping compositional changes constant to evaluate the role played by changes in the returns to labor market attributes. The analysis suggests a distinct role for the education and experience premiums. The secular decline of the experience premium was ubiquitous: it affected the three countries and all education groups, except for the least educated in Argentina. In general, the reductions in the experience premium were stronger among better educated groups. Thus, they are somewhat more important for explaining the reductions of upper-tail (90/50 earnings gap) inequality, although they also had a role in the compression of the lower tail (50/10 earnings gap). In contrast, a falling schooling premium bears a stronger weight in reducing inequality below the median (50/10 earnings gap). This was driven by a much faster decline of the high school premium vis-à-vis workers who had at most completed primary education, compared with the reduction of the college premium. Holding the structure of pay constant, the counterfactual analysis showed that changes in composition, particularly increased education, were inequality enhancing. Hence, while compositional changes may have contributed to increasing wage inequality before the 2000s, they cannot explain the substantial decline in wage inequality of the past decade.

We document important changes in the structure of labor supply during the past two decades. The working-age population is aging and becoming more educated. To link changes in the schooling and experience premiums to the observed changes in labor supply, we built a nested CES model where there is imperfect substitution across experience and education groups. The model allows estimating the elasticity of substitution across experience groups, which ranges from 3.6 among workers with high school completed or less to 5.5 among college-educated workers. The model also yields estimates of the elasticity of substitution across different levels of schooling. The elasticity of substitution across workers with secondary education and high school dropouts is 2.3, while that of college graduates vs. unskilled workers drops to 1.25. These estimates are shown to be very robust to a battery of robustness checks.

The results from the estimation of the model suggest that changes in labor supply have been a significant factor behind the recent evolution of the schooling and experience premiums. The secular aging of the population explains about a half of the decline in the experience premium among unskilled workers, and about a third of the reduction among college graduates. Similarly, the massive increase in the supply of high school graduates goes a long way toward explaining the decline in the high school/primary wage gaps, although in the cases of Argentina and Chile, it misses a mild upward trend during the 1990s. Trends in labor supply do a comparatively worse job in explaining the college premium. A rising supply of college-educated workers cannot account for the increase in the college/unskilled wage gap during the 1990s. During the 2000s, labor supply trends pushed the college premium down, but not enough to explain the full fall, at least in Argentina and Chile.

A key issue for future research is to understand the economic factors behind the movements in relative demand uncovered in the paper. Little is known about the demand for experience, let alone the potential drivers of experience-biased technical change. That the decline in the experience premium was stronger among high school-educated and college-educated workers may suggest a form of skill obsolescence associated with age. The sharp increase in the demand for college educated workers in Argentina and Chile during the 1990s has been the subject of substantial scrutiny, with globalization and technical change as prominent forces (Pavcnik, 2003; Galiani and Sanguinetti, 2003; Bustos, 2011). Less is known about the declining relative demand for college graduates that we find in the three countries during the 2000s.

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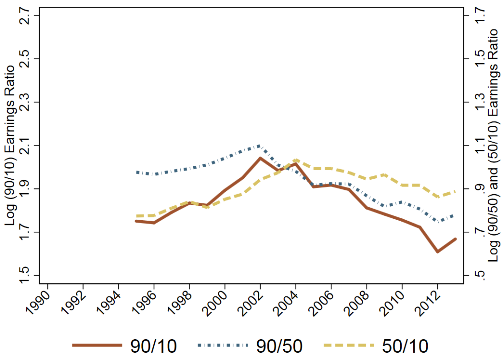
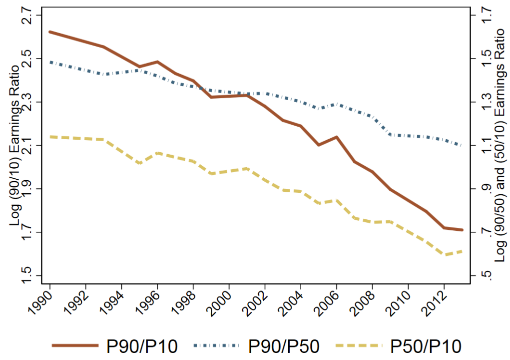
Tables and Figures

Figure 1: Interquantile Log Earnings Ratio by Country

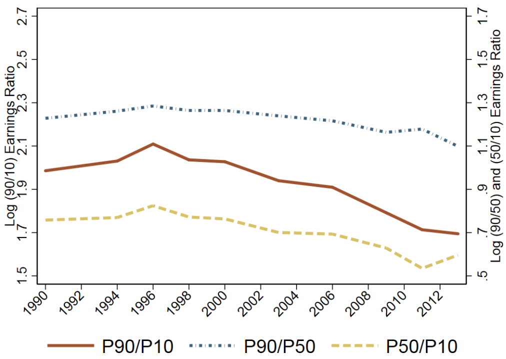
(a) Argentina



(b) Brazil



(c) Chile

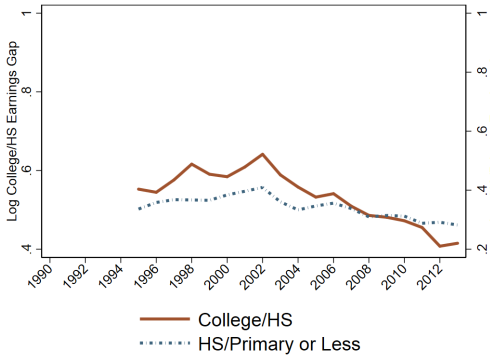


Notes: Sample consists of full-time workers (reported working 35 hours or more) between ages 16 and 65.

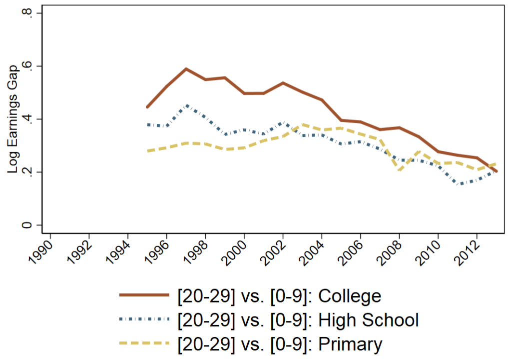
Figure 2: Composition-Adjusted Earnings Gap by Education and Education-Experience

Argentina

(a) Education Premium

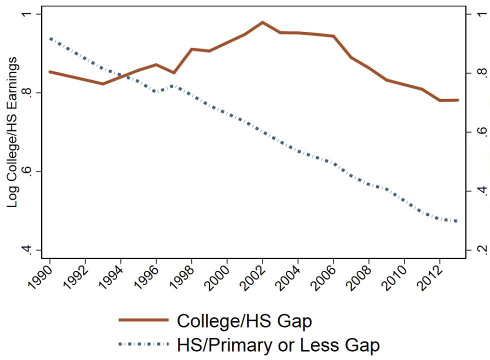


(b) Experience Premium

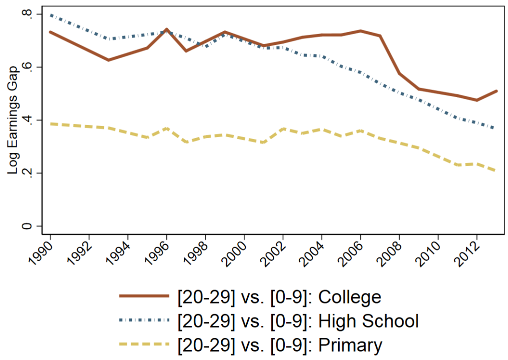


Brazil

(c) Education Premium

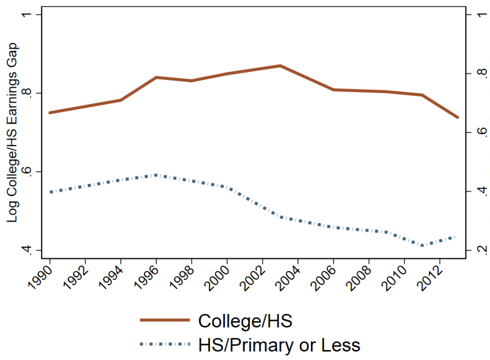


(d) Experience Premium

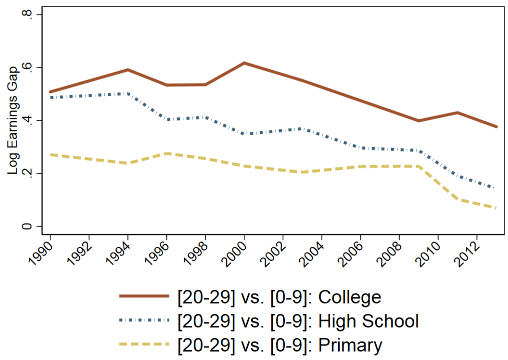


Chile

(e) Education Premium



(f) Experience Premium

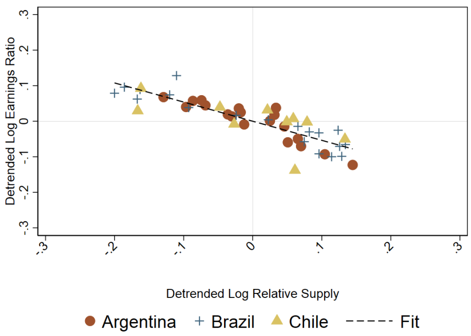


Notes: Sample consists of full-time workers (reported working 35 hours or more) between ages 16 and 65. See Appendix Appendix A.1 for details on the construction of the compositionally adjusted series.

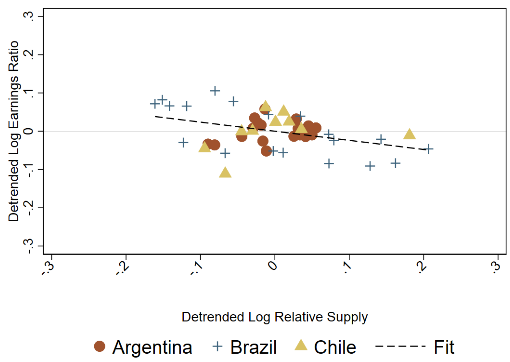
Figure 3: Scatter Plots Premiums vs. Relative Supplies

Education Premium

(a) High School/Primary Gap

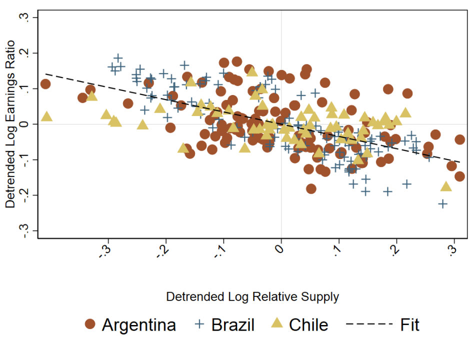


(b) College/High School Gap

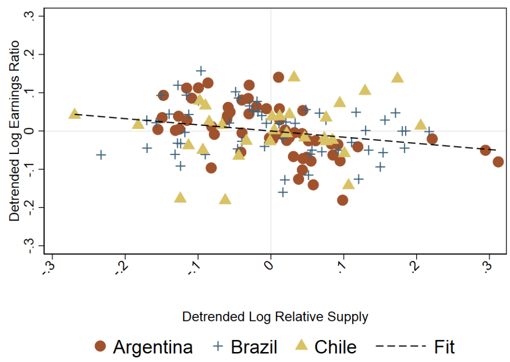


Experience Premiums Conditional on Education

(c) High School– and Primary-educated Workers



(d) College-Educated Workers

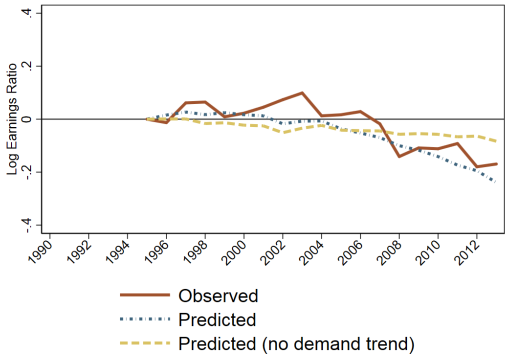


Notes: Detrended log earnings ratio corresponds to the residuals from a regression of observed relative earnings on a linear trend and a country fixed effect. Detrended log relative supplies correspond to the residuals from a regression of observed relative supplies on a linear trend and a country fixed effect. In the case of the experience premiums the residual earnings ratios and relative supplies are calculated separately for the [10-19], [20-29] and [≥ 30] groups relative to the least experienced [0-9] group, and they are all included in the two lower panels.

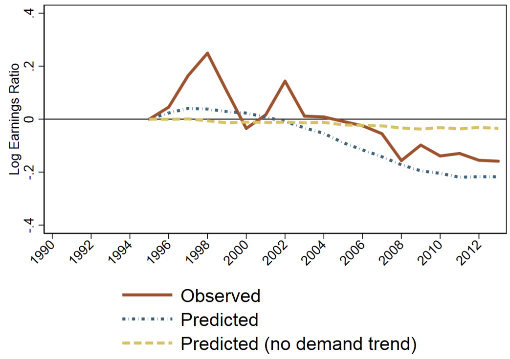
Figure 4: Experience [20-29]/[0-9] Wage Gap: Observed and Predicted Series

Argentina

(a) Unskilled

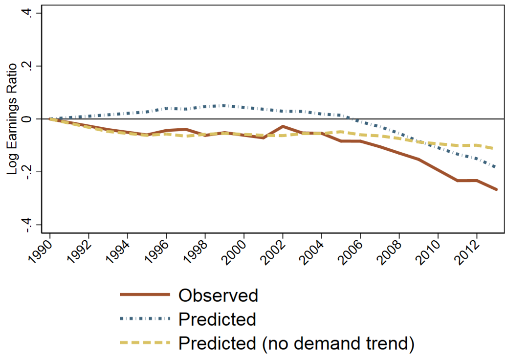


(b) College

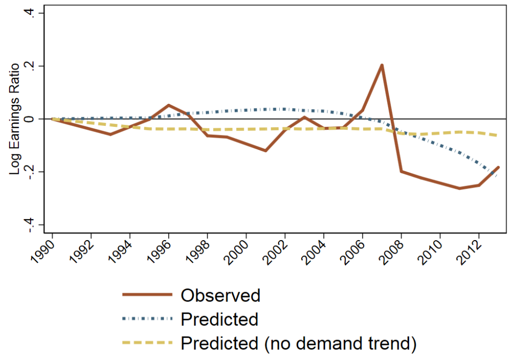


Brazil

(c) Unskilled

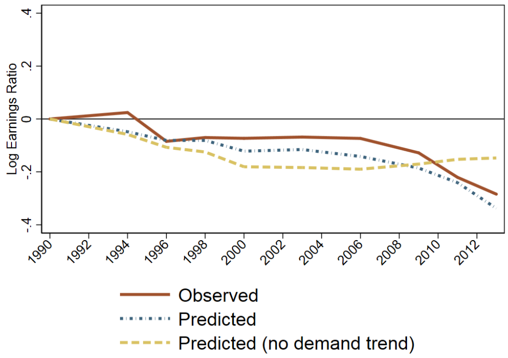


(d) College

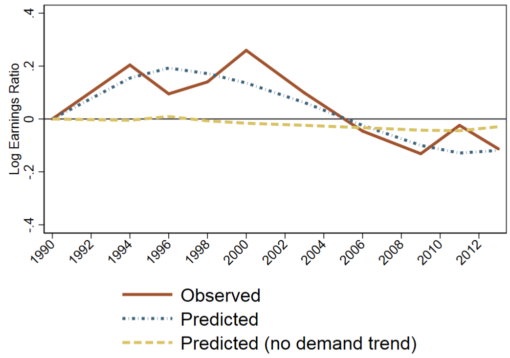


Chile

(e) Unskilled



(f) College

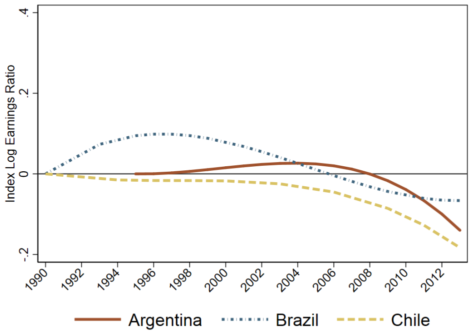


Notes: "Observed" refers to the log ([20-29]/[0-9]) earnings ratio observed in the data for unskilled and college educated workers. "Predicted" refers to the model prediction derived from the estimation of Equations (5.6) and (5.7). The observed and predicted unskilled earnings series are constructed as a weighted average between the primary and high school subgroups, where the weights correspond to the respective labor share. "Predicted (no demand trend)" is the predicted series after substracting the corresponding estimated relative demand.

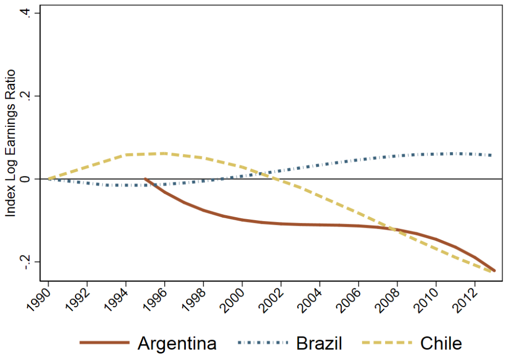
Figure 5: Demand Trends Estimates for Experience Wage Gap Estimation

Potential Experience Groups: [10 – 19]/[0 – 9] Wage Gap

(a) Unskilled: $\log \hat{\phi}_{U2t}$

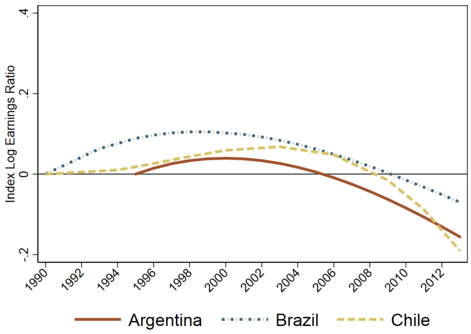


(b) College: $\log \hat{\phi}_{C2t}$

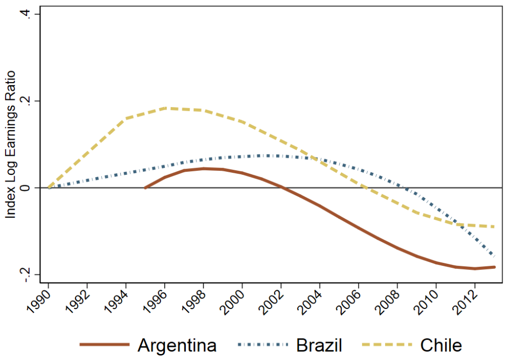


Potential Experience Groups: [20 – 29]/[0 – 9] Wage Gap

(c) Unskilled: $\log \hat{\phi}_{U3t}$

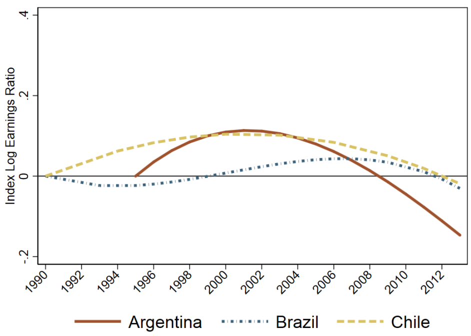


(d) College: $\log \hat{\phi}_{C3t}$

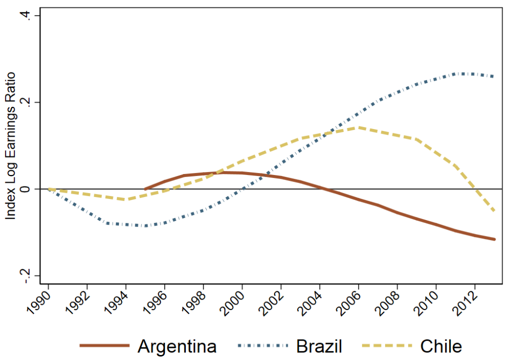


Potential Experience Groups: ≥ 30/[0 – 9] Wage Gap

(e) Unskilled: $\log \hat{\phi}_{U4t}$



(f) College: $\log \hat{\phi}_{C4t}$

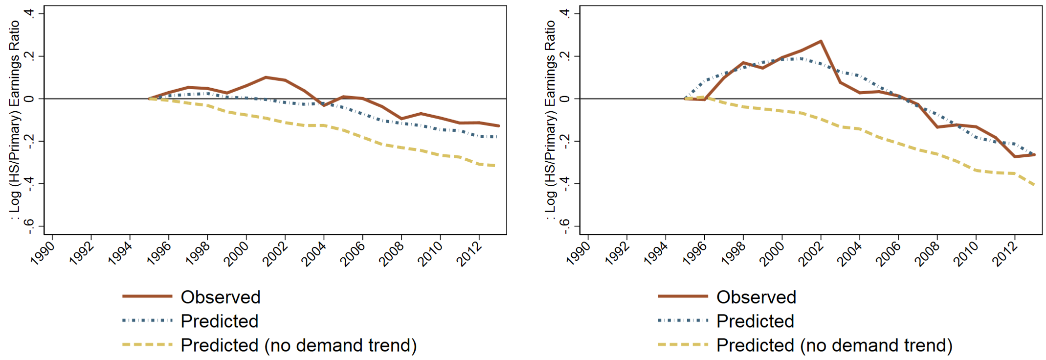


Notes: Relative demand trends between [10-19], [20-29], and [$\geq$ 30] groups relative to the least experienced [0-9] group are country-specific cubic time trends estimated following Equations (5.6) and (5.7). Each series is scaled so that it takes a value of zero in the first year in which data for the country are available.

Figure 6: High School/Primary and College/Unskilled Wage Gaps: Observed and Predicted Series

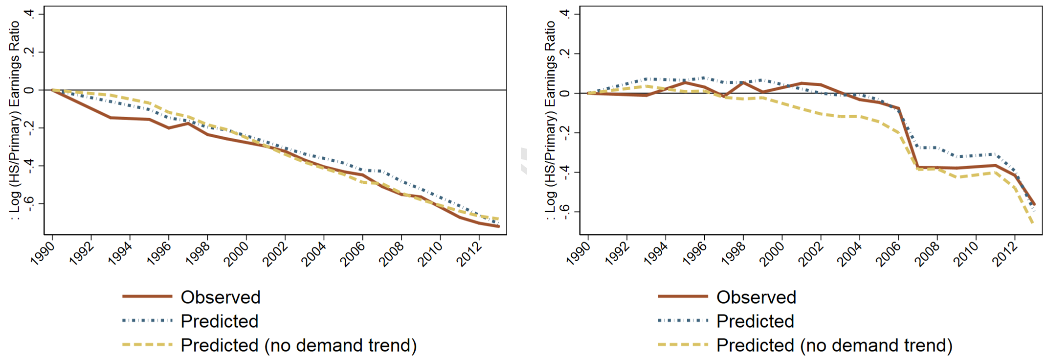
Argentina

(a) High School/Primary (b) College/Unskilled



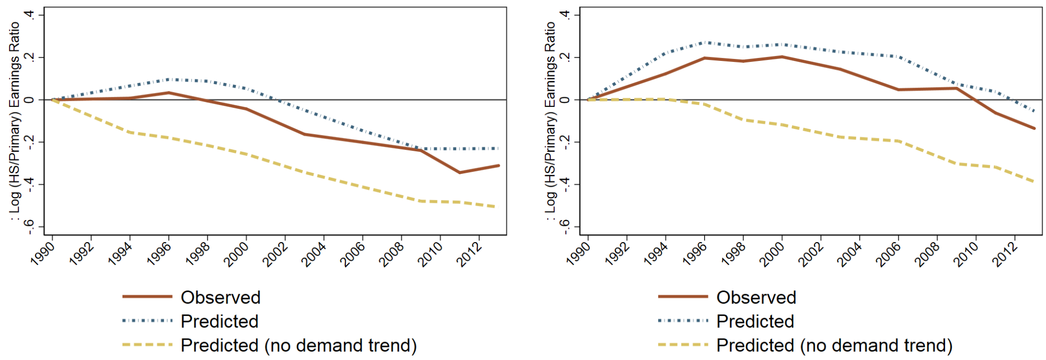
Brazil

(c) High School/Primary (d) College/Unskilled



Chile

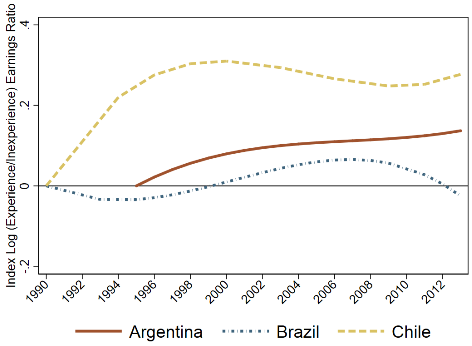
(e) High School/Primary (f) College/Unskilled



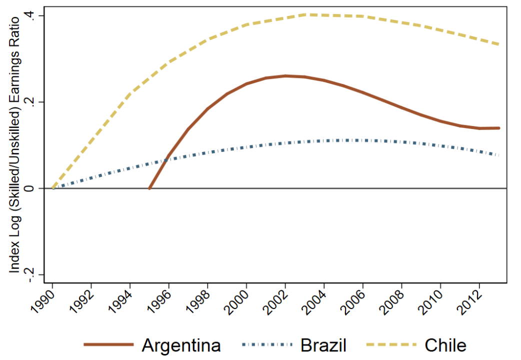
Notes: "Observed" refers to the log (HS/Primary) earnings ratio and log (College/Unskilled) earnings ratio observed in the data. "Predicted" refers to the model prediction derived from the estimation of Equations (5.8) and (5.9). "Predicted (no demand trend)" is the prediction of a modified version the model of Equations (5.8) and (5.9) that omits the respective country-specific time trends. The observed and predicted unskilled earnings series are constructed as a weighted average between the two low-skill subgroups, where the weights correspond to the respective labor share.

Figure 7: Demand Trend Estimates: High School/Primary and College/Unskilled Demand Index

(a) High School/Primary: log $\hat{\beta}_t$



(b) College/Unskilled: log $\hat{\alpha}_t$



Notes: Relative demand trends between high school/primary and college/unskilled are country-specific cubic time trends estimated following Equations (5.8) and (5.9). Each series is scaled so that it takes a value of zero at the first year in which data for the country are available.

Table 1: 100 × Change in Working-Age Population (WAP) and Employment Shares

	Argentina (1995-2013)		Brazil (1990-2013)		Chile (1990-2013)	
	WAP (×100)	Employment (×100)	WAP (×100)	Employment (×100)	WAP (×100)	Employment (×100)
Education						
Primary	-16.60	-16.11	-30.30	-31.54	-24.73	-23.88
High School	8.00	6.88	19.89	19.86	12.22	11.42
College	8.59	9.22	10.41	11.68	12.51	12.46
Pot. Exp.						
[0-9]	0.16	-2.96	1.63	-3.10	-0.11	-6.71
[10-19]	2.78	1.89	-5.64	-4.83	-7.30	-7.73
[20-29]	-0.11	-0.04	0.12	1.45	1.55	2.25
[≥ 30]	-2.67	0.62	1.91	2.55	5.41	8.67
Sex						
Female	-0.33	3.45	0.35	7.04	0.57	10.22

Notes: The working-age population sample includes all individuals between ages 16 and 65. The employment sample includes all individuals between ages 16 and 65 who are employed. Tabulated numbers are changes in the respective shares of each group.

Table 2: Compositional Changes and Inequality Patterns: Oaxaca-Blinder Decomposition Results

	Argentina (1995-2013)		Brazil (1990-2013)		Chile (1990-2013)	
	Est.	[S.E]	Est.	[S.E]	Est.	[S.E]
Log (90/10)						
Overall	-0.091	[0.016]	-0.850	[0.017]	-0.290	[0.021]
Composition	0.056	[0.007]	0.282	[0.016]	0.249	[0.011]
Education	0.054	[0.007]	0.302	[0.016]	0.211	[0.010]
Experience	0.001	[0.001]	-0.003	[0.001]	0.051	[0.002]
Sex	0.002	[0.001]	-0.017	[0.001]	-0.013	[0.001]
Wage Structure	-0.147	[0.018]	-1.132	[0.028]	-0.538	[0.023]
Education	-0.271	[0.113]	-1.153	[0.121]	-1.685	[0.086]
Experience	-0.282	[0.044]	-0.825	[0.095]	-0.497	[0.055]
Sex	-0.049	[0.009]	-0.042	[0.008]	-0.033	[0.007]
Constant	0.454	[0.139]	0.888	[0.230]	1.677	[0.120]
Log (90/50)						
Overall	-0.214	[0.014]	-0.350	[0.011]	-0.149	[0.018]
Composition	0.056	[0.005]	0.222	[0.008]	0.183	[0.009]
Education	0.056	[0.005]	0.229	[0.009]	0.169	[0.009]
Experience	0.001	[0.001]	-0.001	[0.001]	0.030	[0.002]
Sex	-0.001	[0.000]	-0.005	[0.001]	-0.015	[0.001]
Wage Structure	-0.270	[0.015]	-0.572	[0.015]	-0.332	[0.021]
Education	0.084	[0.059]	0.076	[0.041]	-1.021	[0.072]
Experience	-0.253	[0.036]	-0.204	[0.054]	-0.297	[0.053]
Sex	-0.020	[0.005]	-0.013	[0.004]	-0.019	[0.007]
Constant	-0.080	[0.079]	-0.431	[0.090]	1.004	[0.102]
Log (50/10)						
Overall	0.123	[0.018]	-0.500	[0.013]	-0.140	[0.017]
Composition	0.001	[0.005]	0.060	[0.013]	0.065	[0.006]
Education	-0.002	[0.004]	0.073	[0.014]	0.042	[0.006]
Experience	-0.000	[0.001]	-0.002	[0.001]	0.021	[0.001]
Sex	0.003	[0.001]	-0.011	[0.001]	0.002	[0.001]
Wage Structure	0.123	[0.018]	-0.559	[0.024]	-0.206	[0.019]
Education	-0.355	[0.101]	-1.228	[0.095]	-0.664	[0.067]
Experience	-0.029	[0.039]	-0.621	[0.052]	-0.201	[0.027]
Sex	-0.028	[0.008]	-0.029	[0.006]	-0.014	[0.006]
Constant	0.534	[0.124]	1.319	[0.166]	0.672	[0.085]

Notes: Sample consists of full-time workers (reported working 35 hours or more) between ages 16 and 65. Standard errors calculated via bootstrap with 500 replications.

Table 3: Changes in Real Log Hourly Earnings, Employment and Working-Age Population (WAP) across Experience and Education Groups

	Argentina (1995-2013)			Brazil (1990-2013)			Chile (1990-2013)		
	Log Earnings (×100)	(Male) Log Earnings (×100)	Log WAP (×100)	Log Earnings (×100)	(Male) Log Earnings (×100)	Log WAP (×100)	Log Earnings (×100)	(Male) Log Earnings (×100)	Log WAP (×100)
Primary									
(0-9)	28.25	27.66	-10.54	41.36	33.12	-50.03	69.70	66.78	-102.97
	[3.30]	[3.59]	[5.81]	[1.12]	[1.15]	[1.32]	[2.50]	[2.71]	[5.63]
(10-19)	22.16	20.26	29.79	34.88	28.88	-62.61	60.66	62.31	-135.54
	[2.49]	[2.58]	[3.81]	[0.95]	[1.09]	[0.96]	[2.12]	[1.96]	[3.39]
(20-29)	25.23	26.85	50.20	18.41	13.66	-8.90	50.27	49.34	-59.78
	[2.64]	[2.62]	[3.50]	[1.02]	[1.19]	[0.97]	[2.04]	[2.63]	[2.71]
(≥ 30)	18.34	20.14	58.06	26.76	22.70	30.33	39.72	38.29	-2.46
	[2.03]	[2.27]	[1.87]	[0.86]	[1.00]	[0.56]	[1.60]	[1.94]	[1.52]
High School									
(0-9)	25.70	25.81	136.21	-5.72	-16.91	133.55	58.37	56.03	44.27
	[1.93]	[2.37]	[3.30]	[1.15]	[1.48]	[1.68]	[1.63]	[2.19]	[2.25]
(10-19)	19.21	18.90	153.53	-34.97	-42.19	124.68	43.45	40.75	65.68
	[1.95]	[2.39]	[3.53]	[1.28]	[1.51]	[1.56]	[1.83]	[2.31]	[2.93]
(20-29)	8.04	5.16	125.01	-53.35	-54.01	159.44	22.73	25.11	124.87
	[2.49]	[2.91]	[3.73]	[1.99]	[2.38]	[2.18]	[2.34]	[2.95]	[3.32]
(≥ 30)	4.66	4.90	126.75	-51.14	-51.95	187.98	7.74	6.55	145.02
	[3.00]	[3.57]	[3.44]	[3.46]	[3.88]	[2.78]	[3.49]	[4.20]	[3.87]
College									
(0-9)	15.05	10.74	131.90	-39.37	-48.77	132.31	41.27	33.09	99.95
	[2.89]	[3.85]	[4.49]	[1.88]	[2.61]	[2.78]	[3.20]	[4.55]	[4.24]
(10-19)	-4.56	-9.61	158.13	-52.20	-53.35	114.19	30.47	24.00	111.97
	[3.36]	[4.60]	[3.75]	[1.79]	[2.57]	[2.35]	[3.86]	[5.34]	[4.75]
(20-29)	-10.71	-12.70	158.57	-72.88	-75.33	170.38	19.91	17.53	139.55
	[4.25]	[5.38]	[4.83]	[2.73]	[3.74]	[2.98]	[5.15]	[6.15]	[5.33]
(≥ 30)	-17.56	-22.57	201.13	-49.71	-42.66	216.77	16.38	10.61	186.55
	[6.38]	[7.01]	[5.64]	[4.23]	[4.77]	[3.99]	[7.73]	[8.87]	[7.24]

Standard errors calculated via bootstrap with 500 replications. The wage series is calculated using full-time workers (reported working 35 hours or more) between ages 16 and 65. The working-age population series includes all individuals between ages 16 and 65.

Table 4: Model Results: Estimates of the Elasticities of Substitution by Stage

	Demand Trends: Country Specific Polynomials			
	IA	IB	II	III
Rel. Supply Experience: Unskilled	-0.280*** [0.017]		-0.328*** [0.018]	-0.246*** [0.024]
Rel. Supply Experience: College		-0.093 [0.092]		-0.181** [0.086]
Rel. Supply High School/Primary			-0.430*** [0.013]	-0.437*** [0.014]
Rel. Supply College/Unskilled				-0.796*** (0.043)
N	576	288	192	384
R^2	0.895	0.867	0.913	0.956

*** 1 percent ** 5 percent * 10 percent. Standard errors in brackets are calculated via bootstrap with 500 replications, where each bootstrap sample remains fixed for all the stages. Each column presents the results of the estimation of the different stages of the model. Column IA shows the OLS estimates of the inverse of the elasticity of substitution between experienced and inexperienced workers within the unskilled group ($\sigma_{\theta U}$) (see Equation (5.6)); column IB corresponds to the OLS estimates of the inverse of the elasticity of substitution between experienced and inexperienced workers within the skilled group ($\sigma_{\theta S}$) (see Equation (5.7)); column II shows the OLS estimates of the inverse of the elasticity of substitution between the two unskilled subgroups (σ_{δ}), and a second estimate of the inverse of the elasticity of substitution $\sigma_{\theta U}$ (see Equation (5.8)); finally, column III shows the OLS estimates of the inverse of the elasticity of substitution between skilled and unskilled labor (σ_{ρ}), as well as additional estimates from the other elasticities in the model (see Equation (5.9)).

Table 5: Model Results: Robustness Exercises

	Estimated Elast. Step III					
	Baseline	Year FE	Occupied Pop.	Total Hours Worked	Prime Age (25-55)	Inc. Part Time
Rel. Supply Experience: Unskilled	-0.246***	-0.177***	-0.238***	-0.236***	-0.373***	-0.258***
	[0.024]	[0.036]	[0.024]	[0.023]	[0.037]	[0.029]
Rel. Supply Experience: College	-0.181**	-0.159***	-0.191**	-0.202**	-0.192***	-0.148***
	[0.086]	[0.016]	[0.087]	[0.078]	[0.049]	[0.040]
Rel. Supply High School/Primary	-0.437***	-0.398***	-0.428***	-0.421***	-0.424***	-0.459***
	[0.014]	[0.017]	[0.013]	[0.013]	[0.014]	[0.013]
Rel. Supply Skilled/Unskilled	-0.796***	-0.696***	-0.698***	-0.865***	-0.861***	-0.833***
	[0.043]	[0.015]	[0.040]	[0.047]	[0.061]	[0.042]
N	384	384	384	384	384	384
R^2	0.956	0.895	0.940	0.959	0.950	0.961

*** 1 percent ** 5 percent * 10 percent. Standard errors in brackets are calculated via bootstrap with 500 replications, where each bootstrap sample remains fixed for all the stages. Each column presents the results of the estimation of the last stage of the model using a different specification. The first column shows the baseline model. In column two the demand trends are estimated using year fixed effects that are common across countries. The third and fourth columns report the estimates of the elasticities when we use occupied population and total hours worked as measures of labor supply, respectively. In column five, we restrict the sample to include only prime age population (between ages 25 and 55). Finally, the last column reports the results if we include both full-time and part-time workers in the wage series.

Appendix A. Appendix (not for publication)

Appendix A.1. Data and Variable Construction

The household surveys used in Argentina between 1995 and 2003 are waves of the Encuesta Permanente de Hogares (EPH), collected by the Instituto Nacional de Estadística (INDEC). This survey was replaced by the Encuesta Permanente de Hogares Continiua (EPH-C) after 2003, breaking the series. The transition between the EPH and the EPH-C included changes in the questionnaires and the frequency in which the surveys were collected. The geographical coverage in EPH-C was extended to include additional agglomerates. To maintain consistency over the period of study, we only use the agglomerates that are present in both surveys. The EPH and EPH-C are representative for urban areas, but close to 90 percent of the population in Argentina lives in urban centers.

The survey used in Brazil is the Pesquisa Nacional por Amostra de Domicilios (PNAD), collected by the Instituto Brasilero de Geografía y Estadísticas (IBGE). The PNAD is a nationally representative survey that has been carried out on a yearly basis since 1967. We use the different waves starting from 1990. Due to exceptional circumstances, the survey was not collected in 1994 and 2000.

The household survey used for Chile is the Encuesta de Caracterización Socioeconómica Nacional (CASEN). The CASEN is a nationally representative household survey collected by the Ministry of Planning through the Department of Economics at the Universidad de Chile. The survey was first implemented in 1987 and carried out every two years from 1990 to 2000, and every three years thereafter. We use all the waves from 1990 to 2013.

We constructed variables capturing the educational attainment and potential experience of all individuals in the sample. Although the countries we analyze differ in the structure of their educational systems, the SEDLAC project has attempted to homogenize the information from the different countries to make it comparable.[23] We use SEDLAC's coding in the construction of the educational attainment series. In particular, we define five possible levels of educational attainment: i) primary education completed or less; ii) high school incomplete; iii) high school completed; iv) college incomplete; and v) college completed or more. Potential experience is defined as the result of sub-

[23]See CEDLAS and World Bank (2014) for a detailed description of the SEDLAC database.

tracting the total number of years of education completed (plus 6) from the age of the individual.

Although we define five possible levels of educational attainment, we mostly work with three categories: primary or less, high school completed, and college completed. Individuals with incomplete levels of education are distributed equally between the previous and next completed level. For example, mean real hourly earnings of workers with college education are calculated as a weighted average between the observed mean wages of this group and the observed mean wages of workers with college incomplete. The weight of the latter group is equal to half of their actual number. This also implies that in the labor supplies used in the model, the supply of workers with primary education completed or less includes half of the total supply of workers of the high school incomplete category. The supply of workers with high school education completed includes half of the supply of workers with high school incomplete and half of the supply of workers with college incomplete. Finally, the supply of workers with college education completed includes half of the total supply of workers with college incomplete.

Each survey includes a question asking workers for their total monetary income from labor in a reference period. This is the variable that we use throughout the paper to capture labor earnings. The variable is divided by the total number of hours worked to obtain hourly earnings. The series are converted into real terms using the Consumer Price Index of the respective countries.[24] In the main specification, we restrict the sample to individuals between ages 16 and 65, and only use earnings of full-time workers (individuals working 35 hours or more in the reference week).

The composition-adjusted earnings of aggregate groups are constructed using a fixed-weighted average of the different sex-education-experience subgroups. We first run a regression of log hourly earnings on the full set of covariates, which includes indicators for the five education categories, seven dummies for potential experience in five-year intervals, and all possible interactions. The regression is estimated separately for males and females in each available country-year. The predicted log wages from these regressions are evaluated for the 70 subgroups, and a weighted average is estimated when aggregating to broader groups. The weights are equal to the mean employment

[24] Due to inconsistencies found in the official Consumer Price Index in Argentina (see Cavallo (2013)), we use the information from PriceStats (`http://www.statestreet.com/ideas/pricestats.html`) to deflate nominal wages in this country.

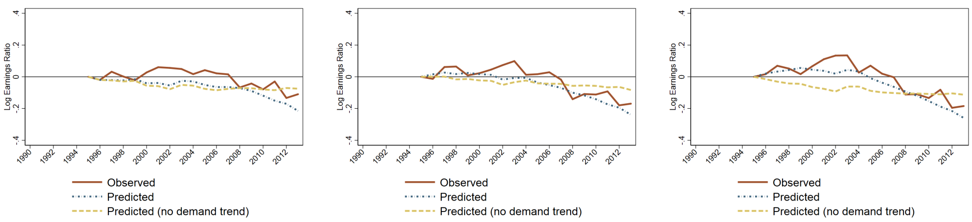
share of each subgroup across all years.

Appendix A.2. Appendix Figures and Tables

Figure Appendix A.1: Observed and Predicted Relative Earnings: All Potential Experience Groups, Unskilled Workers

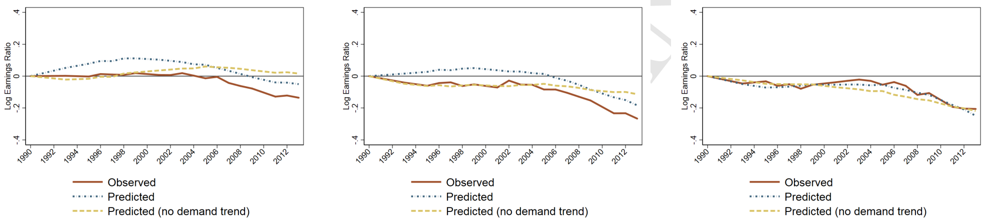
Argentina, Unskilled

(a) [10-19]/[0-9] (b) [20-29]/[0-9] (c) [≥ 30]/[0-9]



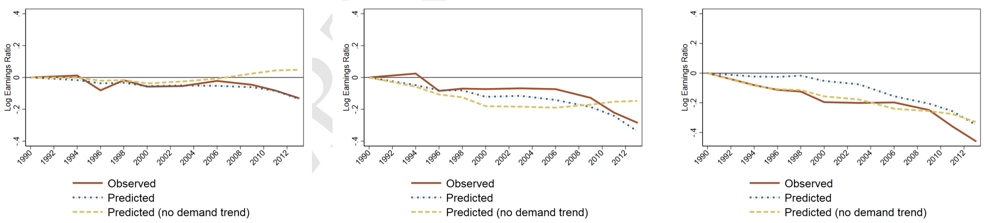
Brazil, Unskilled

(d) [10-19]/[0-9] (e) [20-29]/[0-9] (f) [≥ 30]/[0-9]



Chile, Unskilled

(g) [10-19]/[0-9] (h) [20-29]/[0-9] (i) [≥ 30]/[0-9]

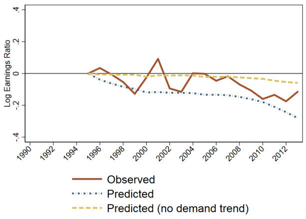


Notes: "Observed" refers to the log earnings ratio between [10-19], [20-29], and [≥ 30] groups relative to the least experienced [0-9] group observed in the data. "Predicted" refers to the respective model prediction derived from the estimation of Equation (5.6). The observed and predicted unskilled earnings series are constructed as a weighted average between the primary and high school subgroups, where the weights correspond to the respective labor share. "Predicted (no demand trend)" is the predicted series after substracting the corresponding estimated relative demand.

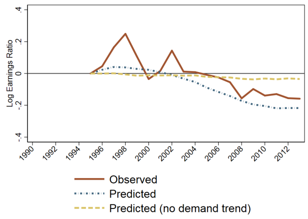
Figure Appendix A.2: Observed and Predicted Relative Earnings: All Potential Experience Groups, College Educated Workers

Argentina, College

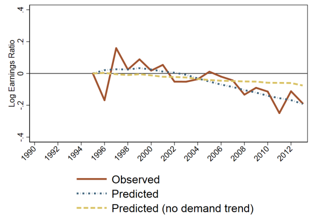
(a) [10-19]/[0-9]



(b) [20-29]/[0-9]

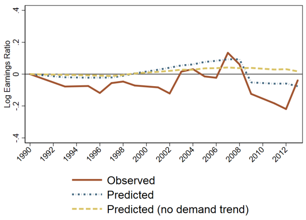


(c) [≥ 30]/[0-9]

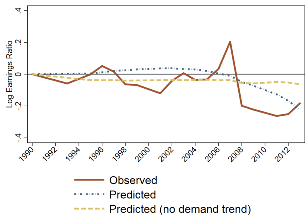


Brazil, College

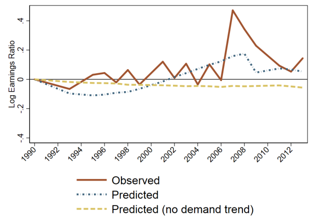
(d) [10-19]/[0-9]



(e) [20-29]/[0-9]

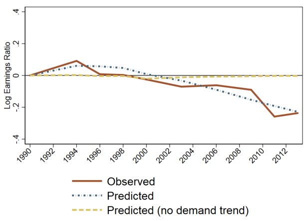


(f) [≥ 30]/[0-9]

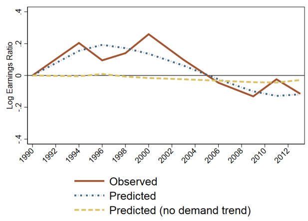


Chile, College

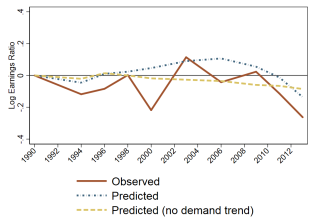
(g) [10-19]/[0-9]



(h) [20-29]/[0-9]



(i) [≥ 30]/[0-9]



Notes: "Observed" refers to the log earnings ratio between [10-19], [20-29], and [≥ 30] groups relative to the least experienced [0-9] group observed in the data. "Predicted" refers to the respective model prediction derived from the estimation of Equation (5.7). "Predicted (no demand trend)" is the predicted series after substracting the corresponding estimated relative demand.

Table Appendix A.1: Changes in Experience Premiums, Employment, and Working-Age Population

	Argentina (1995-2013)			Brazil (1990-2013)			Chile (1990-2013)		
	Log Earnings (×100)	(Male) Log Earnings (×100)	Log WAP (×100)	Log Earnings (×100)	(Male) Log Earnings (×100)	Log WAP (×100)	Log Earnings (×100)	(Male) Log Earnings (×100)	Log WAP (×100)
Primary									
[10-19]/[0-9]	-6.09	-7.40	40.33	-6.48	-4.24	-12.58	-9.04	-4.47	-32.57
	[4.07]	[4.34]	[7.25]	[1.52]	[1.57]	[1.68]	[3.17]	[3.32]	[6.44]
[20-29]/[0-9]	-3.02	-0.81	60.74	-22.95	-19.46	41.13	-19.42	-17.45	43.19
	[4.35]	[4.57]	[6.91]	[1.48]	[1.62]	[1.67]	[3.25]	[3.87]	[6.40]
$\geq$ 30/*[0-9]*	-9.91	-7.52	68.60	-14.60	-10.42	80.36	-29.98	-28.49	100.51
	[3.84]	[4.28]	[6.07]	[1.43]	[1.55]	[1.46]	[2.92]	[3.28]	[5.88]
High School									
[10-19]/[0-9]	-6.49	-6.91	17.31	-29.25	-25.28	-8.87	-14.92	-15.28	21.42
	[2.75]	[3.41]	[4.86]	[1.67]	[2.16]	[2.27]	[2.55]	[3.21]	[3.72]
[20-29]/[0-9]	-17.66	-20.65	-11.21	-47.63	-37.10	25.89	-35.64	-30.92	80.60
	[3.23]	[3.78]	[4.92]	[2.34]	[2.81]	[2.78]	[2.92]	[3.65]	[4.04]
$\geq$ 30/*[0-9]*	-21.04	-20.91	-9.47	-45.43	-35.04	54.43	-50.63	-49.48	100.75
	[3.63]	[4.31]	[4.78]	[3.67]	[4.21]	[3.28]	[3.84]	[4.67]	[4.53]
College									
[10-19]/[0-9]	-19.61	-20.34	26.23	-12.83	-4.58	-18.12	-10.80	-9.09	12.02
	[4.31]	[6.24]	[5.97]	[2.64]	[3.75]	[3.53]	[5.03]	[6.98]	[6.39]
[20-29]/[0-9]	-25.76	-23.44	26.68	-33.50	-26.56	38.07	-21.36	-15.56	39.60
	[5.16]	[6.68]	[6.68]	[3.33]	[4.54]	[4.20]	[5.88]	[7.54]	[6.83]
$\geq$ 30/*[0-9]*	-32.61	-33.31	69.24	-10.34	6.11	84.45	-24.88	-22.48	86.60
	[7.06]	[8.17]	[7.05]	[4.70]	[5.48]	[4.84]	[8.55]	[10.00]	[8.43]

Standard errors calculated via bootstrap with 500 replications. The wage series is calculated using full-time workers (reported working 35 hours or more) between ages 16 and 65. The working age population series includes all individuals between ages 16 and 65. Experience premiums are calculated relative to the group with years of potential experience between zero and nine.

Table Appendix A.2: Changes in Education Premiums, Employment, and Working-Age Population (WAP)

	Argentina (1995-2013)			Brazil (1990-2013)			Chile (1990-2013)		
	Log Earnings (×100)	(Male) Log Earnings (×100)	Log WAP (×100)	Log Earnings (×100)	(Male) Log Earnings (×100)	Log WAP (×100)	Log Earnings (×100)	(Male) Log Earnings (×100)	Log WAP (×100)
HS/Primary									
[0-9]	-2.55 [2.74]	-1.85 [3.05]	146.75 [6.63]	-47.08 [1.42]	-50.03 [1.77]	183.58 [2.15]	-11.33 [2.32]	-10.75 [2.80]	147.24 [6.31]
[10-19]	-2.95 [2.91]	-1.36 [3.10]	123.74 [5.29]	-69.85 [1.58]	-71.07 [1.83]	187.29 [1.90]	-17.21 [2.63]	-21.56 [2.79]	201.22 [4.52]
[20-29]	-17.19 [3.30]	-21.69 [3.57]	74.81 [5.04]	-71.76 [2.28]	-67.67 [2.63]	168.34 [2.47]	-27.54 [2.91]	-24.23 [3.73]	184.65 [4.38]
$\geq$ 30	-13.68 [3.37]	-15.24 [4.09]	68.69 [3.86]	-77.90 [3.60]	-74.64 [3.98]	157.65 [2.89]	-31.98 [3.75]	-31.74 [4.47]	147.48 [4.23]
College/HS									
[0-9]	-10.65 [2.71]	-15.07 [3.56]	-4.32 [5.58]	-33.66 [2.04]	-31.86 [2.73]	-1.24 [3.26]	-17.10 [2.95]	-22.94 [4.10]	55.68 [4.81]
[10-19]	-23.77 [3.66]	-28.50 [4.84]	4.60 [4.99]	-17.23 [2.08]	-11.16 [2.83]	-10.48 [2.82]	-12.98 [4.13]	-16.75 [5.31]	46.29 [5.48]
[20-29]	-18.75 [4.70]	-17.86 [5.91]	33.56 [6.24]	-19.53 [3.23]	-21.32 [4.16]	10.94 [3.65]	-2.82 [5.27]	-7.58 [6.67]	14.68 [6.45]
$\geq$ 30	-22.22 [6.91]	-27.48 [7.86]	74.39 [6.50]	1.43 [5.32]	9.29 [5.84]	28.79 [4.90]	8.65 [8.36]	4.06 [9.47]	41.53 [8.35]

Standard errors calculated via bootstrap with 500 replications. The wage series is calculated using full-time workers (reported working 35 hours or more) between ages 16 and 65. The working age population series includes all individuals between ages 16 and 65.

- A sharp compression of wages in Argentina, Brazil and Chile during the 2000s
- The decline of the schooling and experience premiums are key determinants
- Labor supply trends had an important role in the reductions of the premiums
- Demand for college educated and experienced workers declined during the 2000s.