The Evolution of the College-High School Wage Differential for Males in Brazil: Does an Increasing Supply of College-educated Labor explain it?

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The last two decades exhibit a significant increase in the proportion of workers with a college degree in Brazil. Simultaneously, the college wage premium has been rising for men in the last twenty years. This paper attempts to draw causal relationship between the evolution of college-educated labor supply and the performance of the college wage gap. I develop a model, where the production function has low and high skill labor and no capital. I estimate the elasticity of substitution between college and high school labor types. In addition, I estimate the partial elasticity of substitution between different age groups, for the same school group. I find that a 10% increase in the aggregate college/high school labor supply ratio leads to a 5.2% fall in the wage gap, in the absence of skill-biased technology changes. This can be translated into an elasticity of substitution between college and high school workers of 1.9. The estimate of the partial elasticity of substitution between age groups is very high (around 16). However, such estimate falls to 4.34 when I consider only cohort-specific relative supply of college-type workers. (JEL J31, C23)

I - Introduction[1]

Education attainment of the labor force has mildly increased in the last two decades in Brazil. I am interested in testing if the increase in the number of workers with a college degree has somewhat led to a downward pressure in the college wage premium. The premium is defined as the wage difference between workers with a college degree and workers with a high school degree.

There is recent work that examines the effects of supply changes on the returns to education in the US, Canada and UK. The seminal work, by Lawrence Katz and Kevin Murphy (1992) looks at movements in the college wage premium in the US from 1963 to 1987 and concludes that "it appears to be strongly related to fluctuations in the rate of growth of the supply of college graduates". Their model assumes perfect substitutability among workers in different age groups as long as they have the same level of education. Such an assumption allows them to construct an aggregate education index as a linear combination between the "amount of education" supplied by workers of different ages.

David Card and Thomas Lemieux (NBER,2000) extend Katz & Murphy's model to allow imperfect substitution between workers of different ages. By measuring the impact of changes in the supply of education (of different age groups) on the college wage premium, they are able to estimate the elasticity of substitution between different age groups with the same education level. Once they estimate this parameter, they are able to calculate aggregate indexes for the total supply of high school and college workers in the labor force that take into account the imperfect substitution between age groups. Their estimates for the US suggest that the elasticity of substitution between different age groups is large but finite at 4.4, while the elasticity of substitution between college and high school labor types is in the range of 1.1 to 1.6. Results are similar for both the UK and Canada.

My work examines the applicability of their results in a context of a developing country. In Brazil, the college-high school wage gap for full-time male workers stays fairly flat from 1976 to 1984 at 2.01, then jumps to 2.11 in 1984, and finally jumps once more by

7% in 1988. After 1988, however, it has been fluctuating around 2.20. During the same period, the number of workers with college education in the labor force has being increasing relative to the number of high school workers. This paper examines how changes in the supply of school-related labor has affected the pattern observed in the college-high school wage gap.

I closely follow the two-step estimation method applied by Card & Lemieux (2000). The theoretical model has a production function that uses only labor as input. However, labor can be of two kinds: high school type and college type, combined under a CES technology. It is straightforward to show that the college-high school wage gap for a given age-group depends on both the aggregate relative supply of college labor in period t, and on the age-group specific relative supply of college labor.

To do this work, I create a panel of age/time cells for the period from 1976 to 1998, where each age/time cell contains estimates of college wage premium, college-specific labor supply and high-school-specific labor supply. Each age-specific supply will be measured by the number of individuals in the labor force (weighted by their respective number of hours worked). Aggregate high school labor may not be a linear combination of school labor at different ages, because age groups might be imperfect substitutes. Hence, as a first step, I estimate the elasticity of substitution between age groups within the educational group by regressing the school premium on time dummies, cohort dummies and on the age-group specific relative supply of college labor. Based on the estimated partial elasticity of substitution and estimated cohort-specific productivity factors, I am able to calculate the aggregate college and high school educated labor supply.

The second step consists of estimating the elasticity of substitution between the two school groups by regressing the college-high school wage gap on cohort and age dummies, a linear time trend and on the estimated time series of aggregate relative supply of college labor.

[1] I appreciate the valuable comments of John Karl Scholz, Rodolfo Manuelli and Peter Norman, as well as participants of a seminar at the University of Wisconsin – Madison.

One important issue is how to correctly identify changes in demand for skills from changes in the supply of skills. Following the procedure adopted by Card & Lemieux (2000) and Katz & Murphy (1992), among others, I assume that the effect of skill-biased technical change on college wage premium can be summarized by a time trend. This assumption turns out to be essential for the conclusion that school supply is an important determinant of the college premium.

The main results of the paper are:

a) The estimation procedures reveal that the elasticity of substitution between the two education groups is low and statistically significant. Hence, the estimation indicates that the increasing supply of college workers over time had substantial negative effects on the college wage gap. Depending on the proxy for the age-specific index used, my estimates vary from 1.56 to 1.93 which are close to those observed by Card & Lemieux (2000) for the US.

b) The estimation reveals a very small effect of relative specific college supply in the pattern of wage gap observed by each age group (equivalent to a partial elasticity of substitution across age groups of 16). This might be an effect of lack of specialization in the Brazilian job market. Alternatively, this may be a result of wide changes in the relative supply as an effect of late graduation. Estimating cohort-specific relative supply of college-type labor which are controlled for age effects renders much smaller partial elasticity coefficients, around 4.5.

c) The estimation reveals that the time trend of the college wag gap is significantly positive (between 0.08 and 0.10). This might be an effect of skill-biased technology changes, or accumulation of physical capital, if capital is a complement to skilled labor or a substitute to unskilled labor.

My paper is the only one which actually estimate the elasticity of substitution between age groups in Brazil. There are other studies looking at the impact of labor supply in the school-related wage differential. Fernandes & Meneses (2002) assume three different values for the elasticity of substitution, and estimate the impacts of demand shocks in relative wages. The basic difference between our methods is the identification assumptions. I assume that demand shifts overtime can be translated by a deterministic trend. They take the inverse rout, assuming a given value for the impact of supply on the school premium, and then estimating the impact of demand shifts on wage differentials.

The paper has five additional sections. In section II, the theoretical model is presented. This model is based on Card & Lemieux (2000), and the only difference is that I allow for variations in the age specific productivity factor over time. Section III presents the empirical procedure. Section IV presents the selected sample, and explains how I constructed the panel of wage gaps and school-related labor indexes. In addition, it presents how the constructed series have evolved over time. Section V presents the result of the estimations, when I do not control for the effects of late graduation in the supply of college-type workers. In Section VI, I estimate a cohort-specific relative supply of college-type workers that are not biased by the presence of age effects. Finally, Section VII concludes the paper.

II – Theoretical Framework

Aggregate output depends on two CES sub-aggregates of high school and college labor:

$$C_t = [\sum_j (\alpha_{jt} C_{jt}^{\eta})]^{1/\eta} \tag{1}$$

and

$$H_t = [\sum_j (\beta_{jt} H_{jt}^{\eta})]^{1/\eta}\ , \tag{2}$$

where $-\infty < \eta \leq 1$ is a function of the partial elasticity of substitution σ_A between different age groups j with the same level of education ($\eta = 1 - 1/\sigma_A$). Each age group has specific relative efficiency parameters, α_{jt} and β_{jt}, which vary over time. In other words, those parameters may suffer influence of cohort-specific productivity shocks (e.g. variation in school quality).

In the limiting case of perfect substitutability across age groups, η is equal to 1 and total high-school (or college) labor input is a weighted sum of the quantity of labor supplied by each age group.

I assume that the aggregate production function is also CES:

$$y_t = (\theta_{ct} C_t^{\rho} + \theta_{ht} H_t^{\rho})^{1/\rho}\,, \tag{3}$$

where $-\infty < \rho \leq 1$ is a function of the elasticity of substitution σ_E between the two education groups ($\rho = 1 - 1/\sigma_E$). In this setting, the marginal product of labor for a given age-education group depends on both the group's own supply of labor and the aggregate supply of labor in its education category. Efficient utilization of different skill groups requires that relative wages are equated to relative marginal products[2]:

$$\ln(w_{jt}^H) = \ln(\partial y_t / \partial H_{jt}) = \ln(\theta_{ht} H_t^{\rho-\eta} \Psi_t) + \ln(\beta_{jt}) - 1/\sigma_A \ln H_{jt} \tag{4}$$

$$\ln(w_{jt}^C) = \ln(\partial y_t / \partial C_{jt}) = \ln(\theta_{ct} C_t^{\rho-\eta} \Psi_t) + \ln(\alpha_{jt}) - 1/\sigma_A \ln C_{jt} \tag{5}$$

$$\psi_t = (\theta_{ct} C_t^{\rho} + \theta_{ht} H_t^{\rho})^{1/\rho - 1} \tag{6}$$

Equations (4) and (5) imply that the ratio of the wage rate of college workers in age group j (w_{jt}^c) to the wage of high-school workers in the same age group j (w_{jt}^H) satisfies to the following equation:

$$\ln(w_{jt}^C / w_{jt}^H) = \ln(\theta_{Ct} / \theta_{Ht}) + \ln(\alpha_{jt} / \beta_{jt}) + (1/\sigma_A - 1/\sigma_E) \ln(C_t / H_t) - 1/\sigma_A \ln(C_{jt} / H_{jt})\,. \tag{7}$$

Hence, if the relative employment ratios are exogenous, equation (7) leads to a model for the observed college-high school wage gap r_{jt} of workers in age group j in year t:

$$r_{jt} \equiv \ln(w_{jt}^C / w_{jt}^H) = \ln(\theta_{Ct} / \theta_{Ht}) + \ln(\alpha_{jt} / \beta_{jt}) + (1/\sigma_A - 1/\sigma_E) \ln(C_t / H_t) - (1/\sigma_A) \ln(C_{jt} / H_{jt}) + e_{jt}, \tag{8a}$$

where e_{jt} reflects sampling variation in the measured gap and/or any other sources of variations in age-specific wage premiums[3].

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[2] $\partial y_t / \partial H_{jt} = \partial y_t / \partial H_t \times \partial H_t / \partial H_{jt} = \theta_{ht} H_t^{\rho-1} \Psi_t \times \beta_{jt} H_{jt}^{\eta-1} H_t^{1-\eta} = \theta_{ht} H_t^{\rho-\eta} \Psi_t \times \beta_{jt} H_{jt}^{\eta-1}$.
Similarly, the marginal product of college workers in age group j is $\partial y_t / \partial C_{jt} = \theta_{ct} C_t^{\rho-\eta} \Psi_t \times \alpha_{jt} C_{jt}^{\eta-1}$.

[3] The assumption of exogenous employment ratios is essential for the results. If enrollment decisions are somehow influenced by expected future changes in returns to education, the weighted least squares estimates of the elasticity of substitution will have a positive bias. In other words, the WLS procedure will overestimate the true elasticity of substitution parameters.

According to Equation (8a), the college-high school gap for a given age group depends on both the aggregate relative supply of college labor (C_t / H_t) in period t, and on the age-group specific relative supply of college labor (C_{jt} / H_{jt}). When there is perfect substitution across age groups with the same level of education, the college premium will depend only on the aggregate relative supply of college workers, on the relative technology shock $\theta_{ct} / \theta_{ht}$, and on the age-cohort relative efficiency parameter β_{jt} / α_{jt}.

For purposes of estimation, it is convenient to rearrange Equation (8a) in an alternative form:

$$r_{jt} = \ln(\theta_{Ct}/\theta_{Ht}) + \ln(\alpha_{jt}/\beta_{jt}) - (1/\sigma_E)\ln(C_t/H_t) - (1/\sigma_A)[\ln(C_{jt}/H_{jt}) - \ln(C_t/H_t)] + e_{jt} \qquad (8b).$$

III – Econometric Method

The primary purpose of this paper is to estimate the effect of the aggregate relative supply of college labor on the college-high school wage gap. A problem arises in the attempt to estimate Equation (8a) because aggregate supplies of the two types of labor (C_t and H_t) depend on the elasticity of substitution across age groups, according to Equations (1) and (2). Following Card & Lemieux (2000), a two-step estimation provides a method for identifying both σ_A and σ_E. In the first step, σ_A is estimated from a regression of age-group specific college wage gaps on age group-specific relative supplies of college educated labor, cohort effects (which absorb the relative productivity effect , $\ln(\beta_{jt} / \alpha_{jt})$), and time effects (which absorb the combined relative technology shock and any effect of aggregate relative supply):

$$r_{jt} = b_{t-j} + d_t - (1/\sigma_A)\ln(C_{jt} / H_{jt}) + e_{jt}. \qquad (9)$$

where b_{t-j} are cohort dummies and d_t are time dummies. Given an estimate of $1/\sigma_A$, the relative efficiency parameters α_{jt} and β_{jt} are easily computed by noting that equations (4) and (5) can be transformed into:

$$\ln(w_{jt}^H) + 1/\sigma_A \ln H_{jt} = \ln(\theta_{ht} H_t^{\rho-\eta} \Psi_t) + \ln(\beta_{jt}) \qquad (10)$$

$$\ln(w_{jt}^{C}) + 1/\sigma_A \ln C_{jt} = \ln(\theta_{ct} C_t^{\rho-\eta} \Psi_t) + \ln(\alpha_{jt}), \text{ for all j and t.} \quad (11)$$

The left hand side of these equations can be computed directly using the first step estimate of $1/\sigma_A$. The right hand side can be estimated by a set of time dummies (first term) and cohort dummies (relative efficiency parameters). Given estimates of α_{jt}'s, β_{jt}'s and of η, it is easy to construct estimates of the aggregate supply of college and high school labor in each year (C_t and H_t).

With these estimates in hand, and some assumptions about the time series path of the relative productivity term, $\log(\theta_{ct}/\theta_{ht})$, Equation (8b) can be estimated directly. I follow the previous mentioned literature and assume that $\log(\theta_{ct}/\theta_{ht})$ can be represented as a linear trend.

IV – The College Wage Premium and the Relative Supplies of College Workers by Age

In this section, I present a descriptive overview of trends in the college wage premium for different age groups in Brazil. I also summarize data on the relative supplies of college-educated workers by age group. The data is drawn from the Brazilian Nationwide Household Sample (PNAD), for the years 1976 to 1998[4].

a. The Evolution of the College Wage Premium for Male Workers[5]

Figure 1 shows the evolution of the overall college premium for male workers since 1977. This premium is measured by the coefficient on a regression of log wage on a college dummy and several control variables, for every year in the sample[6]. My estimates

[4] Data is not available for the following years: 1980, 1991 and 1994.

[5] I focus exclusively on the evolution of the college wage gap for men. I believe that this focus is appropriate, given inter-cohort changes in female labor supply that have presumably affected the age profiles of earnings for women in different education groups over the past twenty years. If women in younger cohorts accumulate more actual experience per year of potential experience than older cohorts, this will increase the measured college-high school wage even if the true college premium is fixed. Secular changes in the age profile of the college-high school wage gap may, therefore, be contaminated by these composition effects. Another reason for excluding women from the calculation of the college wage gap is that the sample female hourly labor earnings are very imprecise due to substantial variations in the number of hours worked.

[6] The control variables are: a second degree polynomial for age, five regional dummies and a dummy for self employment. Table A1 in the Appendix shows the results of the regression for each year.

of the college wage premiums are based on the total labor earnings (income from all jobs) of men aged 24 to 58, living in urban areas of Brazil, who are either employed or self – employers and are working at least 40 hours in the week of reference. In addition, extremely low and extremely high hourly wage rates are dropped from the sample[7]. The remaining sample size is 217,206 observations.

The college wage gap has been widening, but not at a constant rate over time. After increasing substantially during the 1980's and reaching 0.82 in 1990, the college-high school differences of the log hourly wage has been oscillating around 0.78 during the 1990's[8]. The college premium increased by 15% from 1977 to 1998, but only 1.2% from 1990 to 1998.

Table 1 presents my estimates of the "college wage premiums" for 5-year age groups, taken at 5-year intervals over my sample period. The estimates are based on differences in mean log average hourly wages between full time workers with a complete or an incomplete college degree (i.e., at least 12 years in school) and those with a complete or an incomplete high school degree (i.e., at least 9 and at most 11 years in school)[9]. The wage gaps are estimated in separate regression models for each cohort in each "year"[10]. Each model includes a dummy for college degree (defined as above), and the following control variables: a linear age term, an indicator of self-employment, five regional

[7] Wages are deflated by the GPI-FGV index, and converted to values of 1996. In 1996 values, the upper bound from 1983 to 1998 is R$ 719.00 per hour, or R$ 28.76 a week (assuming a 40 hour week). This represents respectively the percentile: 98.43%, in 1998; 98.63%, in 1997; 98.25%, in 1996; 98.77%, in 1995; 99% in 1993; 99% in 1992; 99.25% in 1990; 99.25% in 1989; 99.11% in 1988; 99.49% in 1987; 99.44% in 1986; 99.54% in 1985; 100% in 1984 and 1983. The other upper bounds, and respective percentiles in the unrestricted distribution are, in hourly terms: R$ 190.00 (99.7%), in 1982; R$ 177.00 (99.7%), in 1981; R$ 154.00 (99.78%), 1979; R$ 262.00 (99.82%) in 1978; R$ 393.00 (99.81%) in 1977; R$ 971.00 (99.71%), in 1976. The aim of such upper boundaries was to eliminate unrealistic earnings reports. Such misreporting are easily identified, since there is enormous discontinuity above the cutoff points, with the next value being as much as 100,000 larger than the chosen cut-off point. Such misreported data occurred in more than 1.75% of the sample in recent years, with a much smaller fraction for the data from the 1970s. The sample is additionally restricted by the elimination of workers earning less than R$ 6 monetary units per week (25% of the official minimum wage).

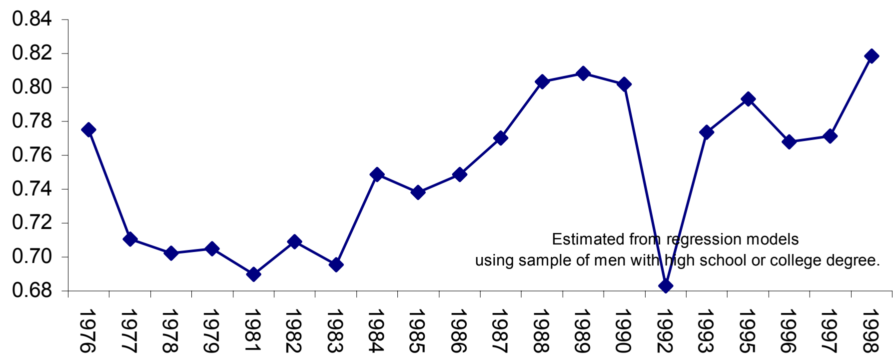
[8] In 1998, a worker in the college group earned on average 2.3 times more than a worker in the high school group.

[9] I can only identify a complete high school or college degree for years after 1992. It is possible that some of the changes in wage can be a result of composition within the education categories, but I do not have how to control for it.

[10] The "year" is a pool of three subsequent years. The only exception is the "year" 1991-93, because there was no survey in 1991.

dummies, and time dummies. The standard error in parenthesis is taken as weights when estimating the model developed in Section II.

Figure 1: College Wage Premium



Comparisons down a column of the table show the changing college premium for a specific age group. Older workers observe a substantial increase in the college premium over time, especially until 1990. Wage gaps increase by 34% from 1977 to 1989 and 1.8% from 1989 to 1998. On the contrary, the wage gap is almost the same for younger workers in 1977 and in 1998. Wage gaps increase by just 2% from 1977 to 1989 and fall by 5% from 1989 to 1998.

Figure 2 follows the age groups 49-53 and 24-28 over the period from 1977 to 1998. The older group represented individuals born between 1923 and 1927, in the beginning of the sample period, and individuals born between 1942 and 1946, in the end of the period. The younger group represents individuals born between 1948 and 1952, in the beginning of the sample period, and individuals born between 1967 and 1971 in the end of the period. If age groups are perfect substitutes, the vertical distance between the two curves may vary either because the return on experience is changing or because the idiosyncrasies of the cohorts.

Table 1: College - High School Wage Differentials - only males							
year/age	24-28	29-33	34-38	39-43	44-48	49-53	54-58
1976-78	0.714	0.790	0.706	0.725	0.665	0.608	0.772
	(.016)	(.018)	(.023)	(.026)	(.031)	(.041)	(.058)
1981-83	0.620	0.725	0.740	0.743	0.703	0.758	0.664
	(.015)	(.015)	(.018)	(.024)	(.029)	(.037)	(.052)
1986-88	0.677	0.783	0.822	0.834	0.740	0.857	0.767
	(.02)	(.019)	(.022)	(.027)	(.037)	(.05)	(.073)
1991-93	0.657	0.691	0.722	0.791	0.796	0.832	0.874
	(.023)	(.024)	(.026)	(.031)	(.040)	(.058)	(.083)
1996-98	0.648	0.767	0.817	0.824	0.808	0.924	0.936
	(.018)	(.018)	(.018)	(.021)	(.026)	(.035)	(.054)

OBS:
(1) College dummy coefficients in a regression model that include as control variables a linear age term, regional dummies, time dummies, and self-employment dummy.
(2) The sample contains a rolling age group. For example, the 24-28 year old group in the 1976-78 sample includes individuals 23-27 in 1976, 24-28 in 1977 and 25-29 in 1978.
(3) College workers are defined as workers who got at least incomplete college degree; High School workers are those who are either high school dropouts or high school graduates (exactly 11 years in school).
(4) Standard Errors in parentheses.

Equation (8a) shows that different evolutions over time of the college wage gap across age groups could happen even in the presence of perfect elasticity of substitution across age groups. Under the assumption of perfect substitutability between age groups, the difference in the college premium between j and j' at time t can be expressed by:

$$r_{jt} - r_{j',t} = \ln(\alpha_{j',t} / \beta_{j',t}) - \ln(\alpha_{j,t} / \beta_{j,t}) + (e_{jt} - e_{j',t}) \qquad (12).$$

Assume a very specific pattern for the age specific productivity factor[11], such that:

$$\ln(\alpha_{jt} / \beta_{jt}) = \varphi_{t-j} + \gamma_j \qquad (13).$$

Then Equation (12) will become:

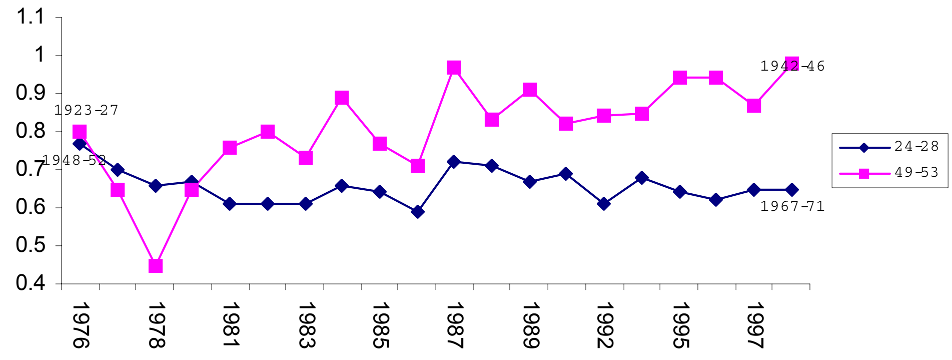
$$r_{jt} - r_{j',t} = (\gamma_{j'} - \gamma_j) + (\varphi_{t-j'} - \varphi_{t-j}) + (e_{jt} - e_{j',t}) \qquad (14).$$

Unfortunately, it is impossible to identify the source of such variation, if changes in the experience returns or changes in cohort-specific "productivity" parameters without strong assumptions. In any case, the variations in $r_{jt} - r_{j',t}$ across age groups are compatible to a

[11] The implicit assumption here is that the returns on experience do not vary over time, and the only source of changes in the premium are attributed to cohort-specific factors.

high elasticity of substitution across age groups[12]. This leads to the second question I try to answer in this paper: Do different evolutions of the school premium for different age groups have anything to do with the evolution of the age specific supply of college labor?

Figure 2: Estimated College-High School Wage Gap for Different Age Groups Over time



b. Relative Supplies

I turn next to an overview of my estimates of the relative supplies of college-educated labor by age group and year. I estimate relative supplies from a broad sample that includes men and women. My estimates of relative supplies of different education groups are based on data for men and women aged 24 to 58, who had worked as an employee or were self-employer at least one hour in the week of reference. The sample size is 1,803,168 observations. To account for differences in the effective supply of labor by different groups, I count the number of hours worked in the week by each worker and weight these hours by the average wage (over all periods) of his (or her) education group.

My main focus is on the evolution of the aggregate college supply over time. Figure 3 shows that there is a clear positive trend on the proportion of workers with more than

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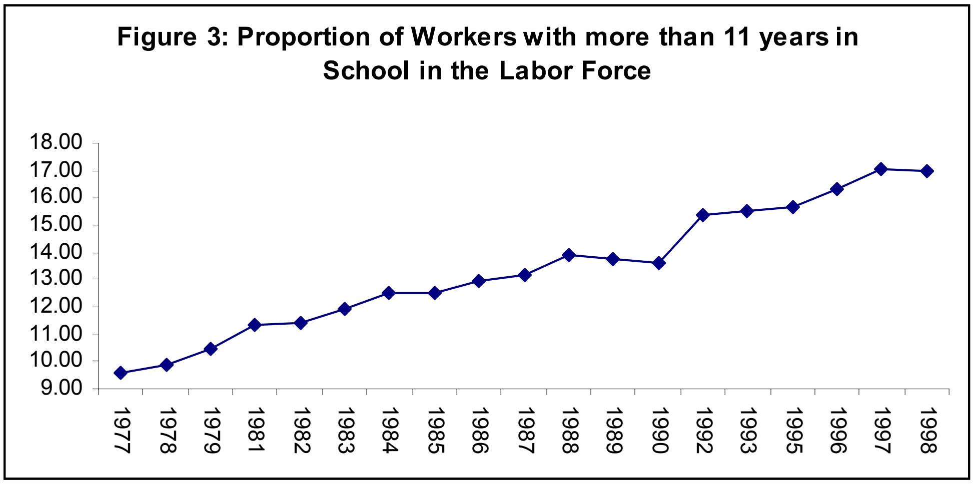
[12] Suppose, for example, that the age-specific "productivity shock" varies across cohorts, in such a way that the college premium is smaller for younger cohorts. One potential explanation for this cohort difference in the college premium could be driven by a sample selection bias. There has been a substantially higher supply of college-level education during the last ten years in Brazil. So, college education has just recently become widely available. However, these new colleges are generally located in poorer areas and offer low quality education. Hence, the college-educated worker of a younger cohort will have, on average, lower skills than the college-educated worker who belong to an older cohort. This will reflected in the larger distance between the college premium of the two birth groups observed in Figure 2, even in the absence of perfect substitution across age groups.

eleven years in school in the labor force, although this proportion is no bigger than 17% at the end of the sample period.



I define the amount of "high school labor" of age group j in year t (H_{jt}) as the total weekly hours[13] worked by high school graduates or dropouts in that age range, plus the total hours of elementary school graduates or dropouts (weighted by their wage relative to the high school group)[14]. The amount of "college labor" of age group j (C_{jt}) is defined as the total number of hours worked by workers with complete or incomplete college degree[15]. Table 2 shows the results of calculations. The estimated relative "college labor supplies", in Table 2C, are used in Section V to estimate the elasticity of substitutions between different age groups.

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[13] I attribute 20 hours worked for individuals who worked less than 40 hours and 40 hours for those working at least 40 hours.

[14] The first group is called "incomplete high school" and the second group is the "elementary school". Both groups are considered in the aggregate labor supply of "high school" workers, with their respective labor efficiency factors. The relative wage of the workers with complete or incomplete elementary school with respect to the wage of "high school workers" is obtained through a regression (for the whole sample period) of the log hourly real wage on a dummy for "high school workers", regional dummies, a dummy for self employment and a squared polynomial of age. The labor efficiency factor of the "elementary school" group is 0.52. Card & Lemieux (2000) use only high school graduates or dropouts in their measure of "high school labor type". In addition, they are able to identify high school dropouts and give respective labor efficiency weights for such group of workers, while I have to consider them with the same weight as those workers who has completed high school.

[15] Card & Lemieux (2000) adopt a similar procedure. The difference is that they can identify college dropouts and then consider the respective labor efficiency factors for that group. In addition, they consider part of the college dropout workers in the high school group, while I have to consider as college-educated workers.

Table 2							
A-School Supply - "High School Labor Supply" by age and year- Men and Women.							
	24-28	29-33	34-38	39-43	44-48	49-53	54-58
1976-78	236,713,531	178,110,812	146,104,657	127,908,086	100,147,384	76,334,866	48,333,458
1981-83	299,082,493	239,629,962	187,009,908	157,651,536	122,209,806	93,682,978	64,463,991
1986-88	380,783,102	319,769,793	259,705,668	203,297,836	156,111,226	115,591,636	78,890,812
1991-93	243,317,801	221,653,991	188,706,706	148,974,792	104,825,954	71,323,342	45,037,681
1996-98	382,365,307	363,455,473	329,210,492	280,303,152	205,388,574	133,125,505	83,187,682
B-School Supply - "College Labor Supply" by age and year- Men and Women.							
	24-28	29-33	34-38	39-43	44-48	49-53	54-58
1976-78	50,810,740	41,364,180	24,814,440	18,517,300	12,684,040	8,245,340	4,714,280
1981-83	66,810,900	65,035,620	46,679,240	29,161,780	18,667,020	12,353,780	7,075,420
1986-88	81,617,120	93,321,780	75,724,380	51,807,620	31,262,900	17,348,500	10,413,900
1991-93	57,458,020	62,228,020	59,588,700	47,979,800	29,843,880	15,832,020	7,899,400
1996-98	88,585,800	105,236,900	104,679,740	93,197,400	68,374,480	39,617,140	17,662,340
C-School Supply - Relative College Labor Supply							
	24-28	29-33	34-38	39-43	44-48	49-53	54-58
1976-78	0.21	0.23	0.17	0.14	0.13	0.11	0.10
1981-83	0.22	0.27	0.25	0.18	0.15	0.13	0.11
1986-88	0.21	0.29	0.29	0.25	0.20	0.15	0.13
1991-93	0.24	0.28	0.32	0.32	0.28	0.22	0.18
1996-98	0.23	0.29	0.32	0.33	0.33	0.30	0.21

OBS:
(1) High School Labor Supply == number of workers with incomplete or complete high school degree (weighted by # weekly hours) + number of workers with elementary school (weighted by # weekly hours and by the relative wage with respect to high school wage).
(2) The average relative wage of the elementary school group (at most 8 years in school) is obtained through a regression of log real wage on age, squared age, regional dummies, and self employment dummy.
(3) College Labor Supply == number of workers with incomplete or complete college (weighted by # weekly hours)
(4) Relative College Labor Supply == college labor supply : high school labor supply.
(5) The number of hours worked is represented by 20 (part time) is the individual worked less than 40 hours and 40 if the individual worked at least 40 hours.

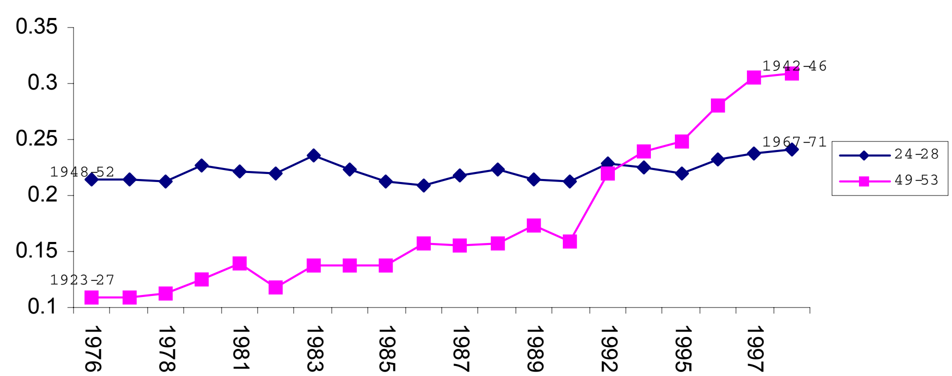
The ratio of college labor to high school labor increases for each age group, but the growth is significantly bigger for the older age groups. Take for example the two age groups in Figure 4 (the same age groups as in Figure 2). The ratio of college to high school labor increases from .21 to .24 from 1977 to 1998 for individuals aged 24-28 years old, while the same ratio increases from .11 to .31 for the workers aged 49-53[16].

If the elasticity of substitution between age groups with the same education degree were low, a supply increase in college workers would lead to a decrease in the wage gap for that specific age group. It is puzzling that exactly the age group that presents the higher increase in relative supply of “college workers” is the one that experiences the higher growth in the college wage premium. A possible explanation for this evidence is that the elasticity of substitution between different age groups is very large. If this is true, the different evolutions of the college wage gap across age groups only can be explained by

[16] The vertical difference between the two curves are not only explained by cohort differences, but by late graduation. Only late graduation can explain that, in 1998, the ratio for the cohorts born from 1942 to 1946 is higher than the ratio for the cohort 1948-52 in 1976. Table 2C shows more clearly the importance of late college graduation in Brazil. For some cohorts (tracked in the diagonals), the college-high school ratio grows up to the age 49-53.

different labor-efficiency parameters for each cohort, like the one modeled in Equations (12) to (14). In this case, the age-specific relative supply of college-educated workers will have no effect on the evolution of age-specific college wage gaps[17].

Figure 4: Relative Supply of College Workers



V – The Effect of Cohort-Specific Supplies on the College Wage Premium of Male Workers[18]

I now turn to the estimation of the effects of the relative supply of college educated workers on the college-high school wage gap. Table 3 presents three estimations of the partial elasticity of substitution. In Column (1), the specification includes time and cohort dummies. In the second specification, the time dummies are replaced by a linear time trend. In the third specification, I use a time trend, age dummies and cohort dummies as controls.

In all three cases, the impact of the age-specific relative supply of college-educated workers on the college wage gap is not statistically different from zero. This seems to

[17]An alternative explanation is that variations in school supply that are driven by late graduation do not affect the college wage gap, while variations in school supply across cohort do affect the wage gap. For example, controlling for age effects, the cohort born in 1967-71 is more educated than the cohort born in 1942-46, and such difference in education will partly explain why the college wage is lower for the younger cohort (see Figure B2 in the Appendix). If this is true, the partial elasticity of substitution between age groups will be much smaller than it seems just by looking Figures 2 and 4. This second interpretation is developed in the Appendix B.

show that the partial elasticity of substitution across age groups is very high. In regressions (1) and (2), the coefficient on the age-specific relative supply is positive.

The addition of age dummies (Column 3) changes the age-specific supply coefficient from positive to negative, although it still not significantly different from zero. This may happen because both the college premium and the college supply ratio increase for initial ages, for different reasons (respectively, positive returns on experience and late college graduation)[19]. Adding age dummies to the RHS in specification (3) has the effect of controlling for this spurious relation across cohorts between college/high school ratios and the wage gaps.

One has to be careful about the ability to identifying the age, cohort and time coefficients. Since Age = Time – Birth Year, the identification of these effects cannot be done without additional assumptions. In specification (1), cohort dummies capture both age effects and cohort-specific fixed effects of the wage gap. In specification (2), the cohort dummies will capture transitory time effects as well (for example, business cycle effects on the college premium), since the time dummies are replaced by a time trend. In specification (3), the life cycle shape of labor earnings is assumed to be fixed over time (captured by the age dummy coefficients), and cohort dummies capture changes in such shape as well as cohort-specific fixed effects and transitory time effects[20].

Because specification (3) has the most complete set of controls and is the only one that shows a negative sign, which is compatible with the theory, I assume $1/\sigma_A$ is -.11 for the

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[18] For all the estimations in this paper, I assume that men and women are perfectly substitutable in the labor force, with identical productivity parameters for a given age and cohort group. This assumption allows me to pool the male and female labor supply and test the impact of it on the college premium of male workers.

[19] Because younger cohorts are observed only for younger ages, they will present a lower relative supply of college-educated workers, compared to cohorts observed at higher ages. At the same time, the positive return on job experience and the fact that younger cohorts are observed only for lower ages makes younger cohorts exactly the ones with a lower college-high school wage gap.

[20] Table A3 in Appendix A shows the complete set of results, including the estimates of age and cohort dummies. The college wage premium reaches the maximum to cohorts born in (1944-48) and then falls continuously. This pattern shows up in all three specifications. Age dummies indicate that college premium is increasing on age, and this should be expected if the returns on experience are systematically higher for college-educated workers than for high school-educated workers. The estimated year effects absorb both the relative technology shock $(\theta_{ct}/\theta_{ht})$ and any effect of changing aggregate supply $((1/\sigma_E - 1/\sigma_A)\ln(C_t/H_t))$. The time effects are small and significant in the specifications (1) and (2), and insignificant in specification (3).

purpose of estimating the cohort-specific efficiency parameters α_{jt} and β_{jt}. Table 4 presents the estimation of equations (10) and (11). The dependent variables are, respectively, the logarithm of the unconditional average hourly real wage of a high school-educated worker w_{jt}^{H} (column (1)) and the logarithm of the unconditional average hourly real wage of a college-educated worker w_{jt}^{C} (column (2)), net of the changes that are related to age-specific school supply. This allows me to estimate the cohort-specific "productivity" factors, or respectively $\ln(\beta_{jt})$ and $\ln(\alpha_{jt})$.

Table 3: Estimated Models for the College - High School Wage Gap, By Cohort and year	(1)	(2)	(3)
Age Group specific relative supply	0.06 (.129)	0.015 (.135)	-0.111 (.268)
Trend		0.044 (.013)	0.033 (.389)
Year Effects:			
1977	-0.087 (.029)		
1982	-0.076 (.024)		
1992	-0.007 (.029)		
1997	0.079 (.036)		
Degrees of Freedom	19	22	17
R- squared	0.85	0.79	0.89

OBS: Standard errors in parentheses.Models are estimated by weight least squares.The dependent variable is the college-high school wage gaps shown in Table 1. Weights are inverse sampling variances of the estimated wage gaps. All models include cohort effects. Model (3) includes age effects as well. The years indicated when reporting the estimated year effects are the mid-points of the year intervals shown in Table 1.

Table 4: Estimated Models for the Log Real Wage of High School and College Workers Equation (10) and (11)	high school	college
Cohort Effects:		
1921	0.74 (.22)	0.895 (.191)
1926	0.904 (.191)	0.911 (.165)
1931	0.813 (.179)	0.901 (.155)
1936	0.659 (.173)	0.847 (.15)
1941	0.685 (.165)	0.852 (.143)
1946	0.596 (.165)	0.861 (.143)
1951	0.501 (.165)	0.717 (.143)
1956	0.402 (.167)	0.569 (.144)
1961	0.287 (.17)	0.396 (.148)
1966	0.186 (.178)	0.276 (.154)
DF	20	20
R-Squared	0.92	0.91

OBS: Standard errors in parentheses. Models are estimated by weight least squares. The dependent variable is the log of the average real wage for each year/age cell. All models include year dummies.

Based on estimates of $1/\sigma_A$, α_{jt} and β_{jt}, I get estimates of the aggregate relative supply index[21], which is plotted in Figure A1. Table 5 presents estimates of the second stage models (based on Equation (8b)) that include both age-group specific relative supplies of college labor, and the aggregate relative supply index. The relative technology shock variable ($\ln(\theta_{ct}/\theta_{ht})$) is assumed to follow a linear trend. All the specifications include cohort and age dummies, as well as a time trend[22].

In Column (1), I use the estimated aggregate supply of college and high school educated workers as the measure of labor supply. A 10% increase in the age-specific relative supply of college-educated workers decreases the college premium by only 0.6%, for that

[21] The evolution of the estimated aggregate relative supply of college workers is plotted in Figure A1.
[22] Controlling for age dummies does not change significantly the coefficient on the aggregate index, but does change the age-specific supply coefficient, which goes from -.007 to -.063.

particular age-group[23]. This implies a very large partial elasticity of substitution of 15.9. but the 5% confidence interval includes infinite elasticity[24].

In contrast, the estimation of equation (8b) shows that the increase in the aggregate supply of college workers substantially depressed the college premium in the last twenty years in Brazil. A 10% increase in the relative supply of college-educated workers drives down the college-high school wage differential by 5.2%, in the absence of non-neutral technology changes. This is compatible to an elasticity of substitution between college and high school workers of 1.93[25].

The time trend coefficient is 0.08 and it is significant at 5%. This means that, absent the age/cohort productivity factor and the changes in school supply, the college wage gap would have increased on average by 8% in each 5-year period. Note that such large, positive and significant coefficients of the time trend contrasts with the trend coefficients in Table 3. The insignificance of the time trend in Table 3 (especially specification (3)) happens because college supply is driving down wage gaps while the skill-biased technology changes are driving up the returns to college education.

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[23] Controlling for age dummies changes the coefficient on the aggregate index from -.44 to -.52, but the age-specific supply coefficient goes from -.007 to -.063.

[24] Variations in the supply ratio across different ages for a given cohort does not seem to depress wage gaps. At the same time, it seems to be big enough to dominate the variations across different cohorts for a given age. In Appendix B, I replace the age-specific relative supply ratio by an estimated supply ratio that is fixed for a given cohort, and re-estimate the elasticity of substitution coefficients. The main conclusion is that, absent the age profile of the college-high school ratio, cohort variations are negatively correlated to the wage gap. The estimated impact of cohort-fixed supply on the college wage gap is -.229 (which is equivalent to a partial elasticity of substitution between age groups of 4.36). The estimate of the elasticity of substitution between school types falls to 1.56.

[25] As noticed before, pooling the male and female labor supply is equivalent to assuming that female labor has the same impact as male labor supply on the male college premium, or that these labor types are perfectly substitutable in the production function. When I adopt the opposite assumption of no substitution, and hence exclude females from the relative supply of college-educated workers, the results are surprisingly different. Both the effects of the aggregate labor supply and the time trend are not statistically significant. This might happen due to the existence of multicollinearity between the time trend and the evolution of the college-high school labor ratios.

Table 5: Models for the College-High School Wage Gap, Equation 8b		
	(1)	(2)
Age-specific Relative Supply	-0.063 (.236)	-0.07 (.234)
Trend	0.083 (.039)	0.094 (.041)
Aggreg. Supply Index for Men and Women	-0.519 (.287)	
Katz-Murphy Aggr. Supply Index		-0.599 (.301)
Degree of Freedom	16	16
R-squared	0.92	0.92
OBS: Standard Errors in parentheses. Models include age and cohort effects.		

Does the estimate of the elasticity of substitution between school groups change if I assume infinite partial elasticity between age groups? In Columns (2), I adopt the simple aggregation of college and high school workers across age, assuming perfect substitution across age-groups, as done in Katz & Murphy (1992). In this case, the estimate of the elasticity of substitution between school groups decreases slightly to 1.67.

VI – Instrumenting the College-High School Ratio

Variations of the supply ratio across different ages for a given cohort does not seem to depress wage gaps. At the same time, it seems to be big enough to dominate the variations across different cohorts for a given age. In this section, I replace the age-specific relative supply ratio by an estimated supply ratio that is fixed for a given cohort, and re-estimate the elasticity of substitution coefficients. The main conclusion is that, absent the age profile of the college-high school ratio, cohort variations are significantly correlated to the wage gap.

I can decompose the relative supply of college-educated workers into two components: an age-specific component (ϕ_j) and a cohort-specific component (λ_{t-j}). This specific decomposition assumes that there is an age profile that is common across different cohorts. The log of the relative supply of college-educated workers will become:

$$\ln(C_{jt} / H_{jt}) = \lambda_{t-j} + \phi_j + u_{jt} \tag{15}$$

Under this decomposition, λ_{t-j} is the projection of the relative supply on the cohort dummies:

$$\lambda_{t-j} = E\left\{ \ln(\frac{C_{jt}}{H_{jt}}) / \phi_j = 0 \right\} \quad (16).$$

Similarly, ϕ_j is the projection of the relative supply on the age dummies:

$$\phi_j = E\left\{ \ln(\frac{C_{jt}}{H_{jt}}) / \lambda_{t-j} = 0 \right\} \quad (17).$$

Table B.1 in the appendix shows the result of a regression of $\ln(C_{jt} / H_{jt})$ in a series of age and cohort dummies in order to find $\hat{\phi}_j$ and $\hat{\lambda}_{t-j}$. All dummies are statistically significant. Figure B1 show plots of $\hat{\phi}_j$. The age profile of the relative supply of college-educated workers has a concave shape, having a very positive slope for initial ages. Figure B.2 shows plots of $\hat{\lambda}_{t-j}$. The cohort profile shows an interesting S-shape. Generations born between 1941 and 1951 experiences the highest <u>increase</u> in the ratio between college and high school educated workers. Nonetheless, $\hat{\lambda}_{t-j}$ reaches the maximum value for the cohort born in 1969-73.

Once I calculate $\hat{\lambda}_{t-j}$, I can use it to estimate the partial elasticity of substitution between age groups, instead of using $\ln(C_{jt} / H_{jt})$. Table 6 presents the estimation of equation (9), modified by the presence of $\hat{\lambda}_{t-j}$:

$$r_{jt} = b_{t-j} + d_t - (1/\sigma_A)\hat{\lambda}_{t-j} + e_{jt} \quad (18).$$

Table 6 can be compared to the results found in Table 3. The estimated partial elasticity of substitution is 4.40 or 4.65, depending if the specification includes or not age dummies. Hence, after filtering the relative supply ratio of the wide variations across age groups, changes across cohorts in the supply of college-educated workers has a negative and significant impact in college wage gap[26].

The next step is estimating equation (10) and (11). I assume $1/\sigma_A = -.227$, and use this value to estimate $\ln(\beta_{jt})$ and $\ln(\alpha_{jt})$, respectively the age/cohort productivity factor of high-school-educated and college-educated and workers. Table 7 is the analog of Table 4 and shows the estimates of $\ln(\beta_{jt})$ (column(1)) and $\ln(\alpha_{jt})$ (column (2)) – the cohort dummies coefficients. I use the estimates of $\ln(\alpha_{jt})$ and $\ln(\beta_{jt})$ to calculate the aggregate supply of college-educated and high-school-educated workers, with the respective cohort efficiency parameters.

Table 8 shows the result of the estimation of Equation (8b), analog to Table 5. Using the cohort-specific measure of relative supply of college-educated workers, instead of the age/time relative supply (as done in Table 5), I find that a 10% increase in cohort supply of the college/high school ratio leads to a 2.3% decrease in cohort-specific college/high school wage gap - when controlling for age and cohort effects[27]. This means that there is a partial elasticity of substitution between age groups of 4.34. The coefficient is statistically significant (the interval does not include zero as a possibility). This confirms my expectation that changes in the supply of college-educated workers across age (related to late graduation) seems to be causing the insignificant coefficients found in Table 3 and Table 5.

[26] When controlling for age effects, for example, a 10% increase in college supply for a given cohort would lead to a 2.3% decrease in college wage gap for that cohort with respect to other cohorts.
[27] The estimate does not change substantially in the specification without age dummies.

Table 6: Estimated Models for the College - High School Wage Gap, by Cohort and year - Equation (9)

	(1)	(2)	(3)
Cohort-specific	-0.219	-0.215	-0.227
relative supply	(.044)	(.048)	(0.04)
Trend		0.045	0.049
		(.007)	(0.006)
Year Effects:			
1977	-0.094		
	(.024)		
1982	-0.079		
	(.022)		
1992	0.001		
	(.022)		
1997	0.092		
	(.024)		
Age Effects:			
29-33			0.034
			(0.023)
34-38			0.008
			(0.022)
39-43			-0.004
			(0.022)
44-48			-0.064
			(0.022)
49-53			0.003
			(0.023)
54-58			
Cohort Effects:			
1924-28	-0.116	-0.131	-0.137
	(.048)	(.051)	(.044)
1929-33	-0.007	-0.017	0.003
	(.041)	(.044)	(.038)
1934-38	0.083	0.065	0.081
	(.037)	(.039)	(.035)
1939-43	0.117	0.101	0.114
	(.035)	(.037)	(.033)
1944-48	0.218	0.202	0.212
	(.04)	(.043)	(.039)
1949-53	0.213	0.195	0.208
	(.044)	(.047)	(.041)
1954-58	0.171	0.151	0.148
	(.047)	(.050)	(.043)
1959-63	0.13	0.114	0.104
	(.047)	(.050)	(.043)
1964-68	0.109	0.086	0.071
	(.051)	(.054)	(.046)
1969-73			
Degrees of Freedom	20	23	18
R- squared	0.85	0.79	0.89

OBS: Standard errors in parentheses.Models are estimated by weight least squares.The dependent variable is the college-high school wage gaps shown in Table 1. Weights are inverse sampling variances of the estimated wage gaps. The years indicated when reporting the estimated year effects are the mid-points of the year intervals shown in Table 1.

Table 7: Estimated Models for the Log Real Wage of High School and College Workers Equation (10) and (11)

	high school	college
Cohort Effects:		
1921	0.459	0.551
	(.24)	(.221)
1926	0.669	0.623
	(.208)	(.191)
1931	0.609	0.661
	(.196)	(.18)
1936	0.48	0.647
	(.189)	(.173)
1941	0.533	0.702
	(.180)	(.165)
1946	0.479	0.779
	(.18)	(.165)
1951	0.42	0.68
	(.18)	(.165)
1956	0.354	0.562
	(.182)	(.167)
1961	0.263	0.399
	(.186)	(.171)
1966	0.178	0.286
	(.194)	(.172)
DF	20	20
R- squared	0.94	0.96

OBS: Standard errors in parentheses.
Models are estimated by weight least squares.
The dependent variable is the log of the average real wage for each year/age cell.
All models include year dummies.

Table 8: Models for the College-High School Wage Gap, By Cohort and Year - Equation (8b)

	(1)	(2)
Cohort-specific	-0.217	-0.229
Relative Supply	(.046)	(.035)
Trend	0.093	0.1
	(.027)	(.021)
Aggreg. Supply Index	-0.604	-0.64
for Men and Women	(.22)	(.201)
Degree of Freedom	22	17
R-squared	0.82	0.92

OBS: Standard Errors in parentheses.
Model (1) includes cohort effects. Model (2) includes age and cohort effects.

The coefficient of the aggregate supply index is equal to -0.64, meaning that a 10% increase in the aggregate relative supply of college-educated workers leads to a 6.4% fall in the college/high school wage gap, which is equivalent to an estimate of 1.56 for the elasticity of substitution between college and high school workers. This is slightly smaller than the one found in Section V[28].

VII - Conclusion

In this paper, I estimate the impact on the college wage of the evolution in the relative supply of college graduate workers in the labor force. I test such evidence in the context of a developing country, Brazil.

My results indicate that the partial elasticity of substitution across age groups (within a given school group) seems to be very large (around 16) when do not controlling for age effects on labor supply. More important, the elasticity of substitution between college-educated and high school-educated labor lies between 1.56 and 1.93. The coefficient of the impact of aggregate labor supply on the college wage gap are all negative and significant. The general conclusion is that aggregate changes in education endowments may be contributing to reduce the huge income inequality in Brazil. This might be happening through the negative impact of a higher contingent in the labor force of college-educated workers on the college wage premium.

One should expect that the substitutability across school-groups should be lower than the substitutability across age groups (for a given school group). College-educated workers are placed in relatively high skill jobs and cannot be easily replaced by high school-educated workers. At the same time, two college-educated workers with different ages can perform the same task with approximately the same efficiency.

Comparing to the early literature, there is no similar study for Brazil, but there are lots of studies looking at the same question for developed countries. Since my technique closely

[28] Also, the presence of a positive (significantly different from zero) long run trend to higher college/high school wage gap is inferred from the linear trend coefficient of .10 in Table B4 (Column (2)). The estimate is close to the one found in Table 5, when the directly observed age-specific supply is used. Hence, the evidence that changes in demand for skills (caused either by trade liberalization or high skill-biased

matches the one used by Card & Lemieux (2000), our results are comparable. They find an elasticity of substitution between the two education groups between 1.1 and 1.6 when they pool men and women together, slightly lower than mine. Their definition of the education group is different from mine, since their high school-educated workers does not include workers with only elementary school. My broader definition of the high school labor equivalent makes the "college labor" and "high school labor" less substitutable. For example, workers with less than five years in school are certainly not replaceable by college-educated workers. So, adopting more strictly-defined education groups could increase my estimate of the elasticity of substitution.

In addition, the partial elasticity of substitution across age-groups that they find for US is around 4.5, which is much lower than the one I find for Brazil when I do not control for age effects on school-related labor supply. Nonetheless, my estimates are close to theirs when I do control for such effects.

Some important qualifications are important to my results. Demand factors may be driving up the demand for "college workers" and consequently the college wage gap. For example, skill-biased technology shocks can increase the demand for college-educated labor, driving up the college wage gap. In contrast, trade liberalization might be increasing the demand for low-skill-labor-intensive goods and decreasing the demand for high-skill-labor-intensive goods, driving downward the college wage gap[29]. Accumulation of physical capital increases the school premium if capital is complement for skilled labor and a substitute for unskilled labor. I assume that such demand-side effects on the wage gap are translated into a linear time trend.

Fernandes & Meneses (2002) adopt the reverse side, taking as given the impact of supply on relative wages, and estimating impact on school premium of shifts in the labor demand . They conclude that demand impacts on wages cannot be translated by a linear trend because it has some sort of structural break in the 90s. They call such break "trade

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technology shocks) is robust to the specification and the variables representing the age specific supply of college workers.

[29] This is the direct application of Stolper-Samuelson Theorem to a developing country, which is abundant in low skill labor. Recent research for Brazil has shown ambiguous evidence of such effect of trade liberalization. For a survey see Moreira & Najberg (1999).

liberalization", and relates it to an increase in the demand for high-school type workers in the 90s.

Another qualification resides in the potential room for endogeneity bias in the results of this paper. Enrollment decisions may be influenced by expected future college premiums. In the presence of such endogeneity, the estimates of elasticity I found will be an upper bound for the true elasticity of substitution parameters. Nonetheless, decisions of enrolling in high school and college study are in general bounded by the lack of schools in Brazil, especially during the period covered in the paper. Hence, I suspect that changes in the employment ratio reveal rather changes in the supply of undergraduate courses (due to deregulation of educational sector) than a result of individual choice to invest more in human capital. If this interpretation is correct, endogeneity will be a minor issue here.

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APPENDIX A – Main Estimations

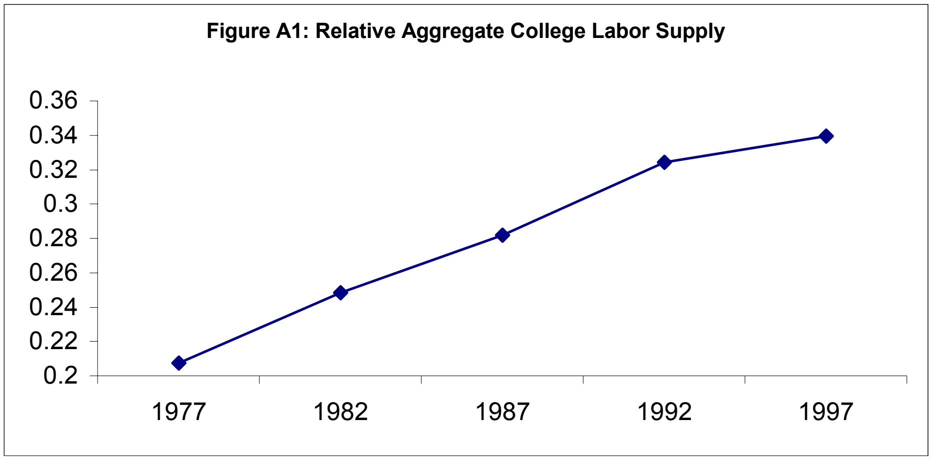
Table A1- Sample Selection - by Birth Year							
year/age	24-28	29-33	34-38	39-43	44-48	49-53	54-58
1976-78	1949-53	1944-48	1939-43	1934-38	1929-33	1924-28	1919-23
1981-83	1954-58	1949-53	1944-48	1939-43	1934-38	1929-33	1924-28
1986-88	1959-63	1954-58	1949-53	1944-48	1939-43	1934-38	1929-33
1991-93	1964-68	1959-63	1954-58	1949-53	1944-48	1939-43	1934-38
1996-98	1969-73	1964-68	1959-63	1954-58	1949-53	1944-48	1939-43

TABLE A2: Dependent Variable: Log real hourly wage										
	1976	1977	1978	1979	1981	1982	1983	1984	1985	1986
Age	0.146	0.144	0.129	0.145	0.142	0.151	0.164	0.162	0.164	0.153
Age^2	-0.002	-0.002	-0.001	-0.002	-0.001	-0.002	-0.002	-0.002	-0.002	-0.002
SE	0.117	0.149	0.158	0.165	0.08	0.024*	0.056	-0.003*	-0.013*	0.057
NE	-0.161	-0.175	-0.187	-0.142	-0.17	-0.209	-0.13	-0.183	-0.023	-0.188
DF	0.239	0.321	0.284	0.262	0.203	0.191	0.225	0.186	0.193	0.139
NO	-0.07*	-0.081*	-0.035*	-0.126	-0.155	-0.117	-0.092	-0.075*	-0.018*	-0.09
CW	-0.052*	-0.051*	-0.031*	-0.077*	-0.102	-0.135	-0.08	-0.098	-0.061*	0.026*
SELF	-0.078	-0.053	-0.14	-0.089	-0.282	-0.271	-0.247	-0.198	-0.184	-0.017*
COL.	0.775	0.71	0.702	0.705	0.69	0.709	0.695	0.749	0.738	0.749
	1987	1988	1989	1990	1992	1993	1995	1996	1997	1998
Age	0.142	0.162	0.146	0.103	0.133	0.107	0.101	0.107	0.101	0.091
Age^2	-0.001	-0.002	-0.002	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001
SE	0.026*	0.062	-0.023*	-0.014*	0.027*	-0.02*	0.067	0.009*	0.032*	0.025*
NE	-0.205	-0.242	-0.322	-0.293	-0.352	-0.326	-0.37	-0.381	-0.376	-0.341
DF	0.106	0.162	0.206	0.224	0.121	0.347	0.29	0.293	0.318	0.337
NO	-0.141	-0.141	-0.124	-0.034*	-0.271	-0.207	-0.163	-0.189	-0.199	-0.209
CW	-0.009*	-0.088*	-0.083*	-0.017*	-0.198	-0.152	-0.129	-0.202	-0.187	-0.191
SELF	-0.118	-0.252	-0.066	-0.041*	-0.263	-0.186	-0.123	-0.089	-0.074	-0.195
COL.	0.77	0.803	0.808	0.802	0.683	0.773	0.793	0.768	0.771	0.818
* Not significant at 5%										

Table A3: Estimated Models for the College - High School Wage Gap, by Cohort and year

	(1)	(2)	(3)
Age Group specific relative supply	0.06	0.015	-0.111
	(.129)	(.135)	(.268)
Trend		0.044	0.033
		(.013)	(.389)
Year Effects:			
1977	-0.087		
	(.029)		
1982	-0.076		
	(.024)		
1992	-0.007		
	(.029)		
1997	0.079		
	(.036)		
Age Effects:			
29-33			0.078
			(.038)
34-38			0.078
			(.03)
39-43			0.089
			(.03)
44-48			0.054
			(.041)
49-53			0.143
			(.062)
54-58			0.160
			(.087)
Cohort Effects:			
1924-28	-0.148	-0.159	-0.134
	(0.053)	(0.056)	(0.044)
1929-33	-0.092	-0.09	0.003
	(0.058)	(0.059)	(0.038)
1934-38	-0.054	-0.055	0.081
	(0.067)	(0.069)	(0.035)
1939-43	-0.091	-0.079	0.114
	(0.086)	(0.089)	(0.033)
1944-48	-0.083	-0.058	0.212
	(0.119)	(0.122)	(0.039)
1949-53	-0.132	-0.104	0.208
	(0.128)	(0.132)	(0.041)
1954-58	-0.186	-0.161	0.148
	(0.126)	(0.129)	(0.043)
1959-63	-0.215	-0.191	0.104
	(0.113)	(0.116)	(0.043)
1964-68	-0.242	-0.229	0.071
	(0.106)	(0.111)	(0.046)
1969-73	-0.342	-0.312	
	(0.095)	(0.1)	
Degrees of Freedom	19	22	17
R- squared	0.85	0.79	0.89

OBS: Standard errors in parentheses.Models are estimated by weight least squares.The dependent variable is the college-high school wage gaps shown in Table 1. Weights are inverse sampling variances of the estimated wage gaps. The years indicated when reporting the estimated year effects are the mid-points of the year intervals shown in Table 1.



Appendix B – Instrumenting the College-High School Ratio

Table B1: Estimated Models for the Relative Supply of College-Educated Workers	
Age Effects:	
29-33	0.249
	(.022)
34-38	0.334
	(.024)
39-43	0.395
	(.026)
44-48	0.47
	(.027)
49-53	0.519
	(.03)
54-58	0.551
	(.032)
Cohort Effects:	
1924-28	0.126
	(.042)
1929-33	0.326
	(.04)
1934-38	0.533
	(.04)
1939-43	0.8
	(.039)
1944-48	1.148
	(.042)
1949-53	1.327
	(.043)
1954-58	1.388
	(.044)
1959-63	1.365
	(.046)
1964-68	1.412
	(.049)
1969-73	1.416
	(.056)

OBS: Standard errors in parentheses.Models are estimated by weight least squares.The dependent variable is the college-high school wage gaps shown in Table 1. Weights are inverse sampling variances of the estimated wage gaps.