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Should I Stay or Should I Go? Career Choices for Young Workers in Latin America*

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Abstract

Labor market institutions shape the return to workers' skills. They define the incentives of firms and workers to invest in general and specific skills, affecting the returns to experience and tenure. This paper presents an empirical assessment of this hypothesis. We take advantage of rich administrative data from Brazil and Chile, two countries with different regulatory frameworks. We focus on young workers as they face to a greater extent the trade-off between accumulating general or specific skills. The analysis is motivated by a simple conceptual framework, from which we derive empirical predictions. It takes into account the potential endogeneity of occupational choices, while allowing for heterogenous returns to tenure and experience. We find positive average returns to both dimensions in both countries, with larger returns to tenure in Brazil and larger returns to experience in Chile. This is consistent with the original conjecture as the more rigid labor market legislation in Brazil encourages firms to carry out additional investments in young workers's specific skills. Chile's institutions, on the contrary, promote the acquisition of general skills. We further examine how these returns differ by educational attainment, an important policy consideration in the region, and find that more educated workers in both countries face larger returns to experience and tenure relative to their less educated counterparts.

JEL classification: J01, J24, J31, J60

Key words: Labor markets, youth employment, returns to skills, occupational mobility, heterogeneity

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1 Introduction

Labor market regulations affect how human capital is rewarded across different institutional frameworks. In particular, certain regulations may change the returns to general and specific skills, by affecting workers' mobility decisions and firms' incentives to train their workers (Acemoglu and Pischke, 1999). As a result, as governments across the world implement policies aimed at improving workers' outcomes, gaining a better understanding of the interaction between regulations and the returns to skill is of particular interest (Vezza, 2014; Carreño *et al.*, 2016). In fact, policies such as subsidized internships and job training programs are often focused on boosting young workers' skills, for whom the early part of their career is of crucial importance, as the first years in the labor force account for a major share of lifetime wage growth (Murphy and Welch, 1990; Topel and Ward, 1992). Nonetheless, while these policies may have heterogeneous impacts across different regulatory frameworks, there is limited empirical evidence on the interaction between regulations and the returns to different types of skill for young workers.

In this paper, we examine the returns to experience and tenure in two Latin American countries with vastly different labor market institutions, Brazil and Chile. Chile's labor market can be described as flexible, with low firing and hiring costs, while Brazil has extensive hiring and firing restrictions, various limits on short-term contracts and a large minimum wage (Heckman and Pagés, 2000; Corseuil *et al.*, 2013). [1] While an extensive theoretical literature has argued that increasing experience and tenure should result in higher wages (Becker, 1964; Becker and Stigler, 1974; Lazear, 1981), the empirical evidence on this topic has largely focused on developed countries, partly due to data limitations. To overcome this problem, we take advantage of administrative matched employer-employee data for both Chile and Brazil and examine the labor market outcomes of young workers from 2003 through 2012. Our data includes information on wages, labor market transitions, participation decisions, labor market experience, different measures of tenure, age, gender, and educational attainment, allowing us to track young workers' labor market trajectories and capture accurate information on earnings.

This comparison is crucial to analyze how markets reward different types of human capital

[1] In April of 2016, Chile's congress passed a labor reform to strengthen unions. The new legislation made it more difficult for businesses to replace striking workers, and companies were prohibited from extending benefits to non-unionized employees. Given the period of time covered in this study, these changes do not affect our analysis.

across distinct institutional frameworks. Theoretically, as long as experience reflects general skills and tenure specific ones, labor markets with larger frictions should exhibit greater investment in both types of skills (Acemoglu and Pischke, 1999). By comparing a highly regulated labor market to a less-regulated one, we are able to shed light on the relative importance of experience and tenure for young workers' wage growth, a first in the empirical literature.

To estimate these returns, we follow the empirical strategy introduced by Dustmann and Meghir (2005). This procedure identifies the returns to experience and tenure using a sample of workers who have been involuntarily displaced from their jobs, effectively exploiting exogenous changes in tenure. To account for the endogeneity of post-displacement labor force participation, we implement a control function approach, which allows us to estimate heterogeneous returns to experience and tenure. We extend their framework by including workers displaced in mass layoffs, defined as events in which the firm lays off at least 50 percent of its workforce. We find large and positive returns to experience in both countries, and higher returns to tenure in Brazil than in Chile, a result which may be driven by Brazilian firms carrying out additional investments in their workers' specific skills, given the country's regulated labor market. We extend our analysis by examining heterogeneous returns by workers' educational attainment and find larger returns to experience for more educated workers in both Brazil and Chile. Lastly, we compare our results to those found in the developed country literature and highlight important similarities and differences. Since the estimated returns to experience and tenure are different across the two countries, our results highlight the need for policymakers to design policies that adequately address the needs of young workers in their context, as one size will not fit all.

The paper is organized as follows. Section 2 discusses the relevant features of labor markets in Brazil and Chile. Section 3 presents a conceptual framework which analyzes how returns can vary with accumulation of post-schooling skills. Section 4 introduces our empirical strategy, describes our data sources and presents summary statistics. Section 5 presents our empirical results for all workers and Section 6 decomposes the return by educational attainment. We discuss our results and conclude in Section 7.

2 Labor Market Institutions

Given the prevalence of features such as large minimum wages, strong unions and high severance costs, labor markets in Latin America have often been characterized as highly regulated (Edwards and Lustig, 1997). Nevertheless, the nature of these provisions varies significantly across countries in the region. For instance, Heckman *et al.* (2004) argue that Chile has one of the lowest total costs of compliance with employment protection laws in the region and Kaplan (2009) found that Chile has the most flexible labor market among fourteen Latin American countries.

Chile's labor market structure largely stems from pro-market reforms implemented during the mid-1970s, which included strong restrictions to union activity and the reduction of fines for dismissal (Montenegro *et al.*, 2004). While the return to democracy in the early 1990s restored some labor protections, including an increase in the maximum severance payment from five to eleven months of salary and re-introducing the need to firms to provide proof for dismissals with cause, labor market regulations have remained lax through our period of analysis. Furthermore, the structure governing collective bargaining agreements did not undergo significant changes and workers' membership in unions remains voluntary. Lastly, the most recent data from the 2017 *Doing Business* project shows that Chile has a ratio of the minimum wage to value added per worker of 0.20, the lowest value in South America.

Brazil's labor market institutions, on the other hand, have been described as rigid and inflexible. For instance, Heckman *et al.* (2004) showed that Brazil has the highest costs of compliance with employment protection laws in Latin America, except for Uruguay. Moreover, unions play a major role in the formal sector, as union dues are mandatory for workers, resulting in all formal sector workers being represented by a trade union (De Barros and Corseuil, 2004). We note that while all formal sector workers are effectively covered by unions, 21 percent of workers are active union members, which far exceeds Chile's unonization rate of 11.5 percent (Bensusán, 2016). Furthermore, while 60 percent of Brazilian workers are subject to collective bargaining agreements, only 9.5 percent of workers in Chile engage in collective bargaining. Brazil also requires firms to pay a large monetary compensation for dismissed employees, equivalent to 40 percent of accumulated funds in a worker's individual savings account in cases of dismissals without just cause. Similarly, the minimum wage has increased significantly in recent years: growing by a total of 119 percent in real

terms between 1996 and 2012. As a result, the Kaitz Index, indicating the ratio of the minimum wage to the average wage, has increased from 0.44 in 2003 to 0.52 by 2012 in Brazil, while in Chile, it has remained flat around 0.41-0.44 (Maurizio and Vazquez, 2016).

In short, as it is costlier for Brazilian firms to hire and fire workers, and these firms face a larger minimum wage and stronger unions than in Chile, Brazil's labor market can be regarded as more frictional than Chile's. We next introduce a conceptual framework which allows us to understand how differential labor market regulations affect the accumulation of post-schooling skills and how these differences should theoretically impact the estimated returns to experience and tenure in these two countries.

3 Methods

3.1 Conceptual Framework

There is a long-standing literature explaining why accumulating experience and tenure results in wage growth. The mechanisms proposed go from improved labor matches for more experienced workers (Jovanovic, 1979) to deferred compensation packages for longer-tenured workers (Lazear, 1981). And, of course, Becker (1964) and Mincer (1974) highlighted the role of post-schooling skill acquisition for explaining the association between wage growth and accumulated experience and tenure. In this context, this literature has pointed out the important differences between general skills, which increase a worker's marginal productivity in all firms, and specific skills, which only increase productivity in the employing firm (for early insights into the subject, see Becker, 1964).

This section presents a simple conceptual framework of a worker's decision to acquire skills after entering the labor force, and show how such skills can explain the relationship between experience/tenure and wage growth. Our conceptual model is inspired by the early literature on post-schooling skill accumulation (Mincer, 1974), which we extend to separately allow for accumulation of general and/or specific skills.

Consider a worker i with entering the labor force in year 0, for whom P_0 represents her potential initial earnings capacity after completing her formal education.[2] In this model, we allow workers to make two types of investment in post-schooling skills in each period t in which they participate

[2]Unlike in Mincer (1974), we do not formally model the schooling decision nor estimate the returns to education.

in the labor force. Their investment decisions can either focus on general skills (k_t), which increase productivity at all firms, or specific ones (d_t), which are only useful if the worker remains employed at the training firm. We express skill-investment costs as a share of potential earnings in period t, P_t, which include both direct costs as well as the opportunity costs of time devoted to skill acquisition. Investment costs in general skills (C_t) and in specific skills (S_t), are thus given by $C_t = k_t \times P_t$ and $S_t = d_t \times P_t$.

Letting ρ_t represent the average return to general skill investments made in time period t and η_t the return to specific skills, we can express potential earnings at time t as:

$$P_t \equiv P_{t-1}(1 + \rho_{t-1}k_{t-1} + \eta_{t-1}d_{t-1}). \tag{1}$$

Assuming the rate of return to general and specific skills is constant across years and equals ρ_0 and η_0, respectively, we can write equation (1) as follows:

$$\ln P_t \approx \ln P_0 + \rho_0 \sum_{j=0}^{t-1} k_j + \eta_0 \sum_{j=\tau}^{t-1} d_j$$

where τ represents the first period of employment at the current firm. Finally, since observed earnings are defined as potential earnings minus investment costs, without further assumptions, we can express the relationship between individual i's earnings in period t, experience $\text{Exp}_i t$ and tenure $\text{Ten}_i t$ as follows:

$$\ln Y_{it} = \ln P_0 + g(\text{Exp}_{it}; \rho) + f(\text{Ten}_{it}; \eta) + \xi_{it} \tag{2}$$

where ξ_{it} is an idiosyncratic earnings shock, and $g(\cdot; \rho)$ and $f(\cdot; \eta)$ are general functions of experience and tenure with associated parameters ρ and η, respectively. These parameters capture the returns to experience and tenure and may include a general component as well as a person-specific component, to account for heterogeneous returns to human capital. Furthermore, the relationship between investment in specific skills and tenure may be not necessarily be linear, as both the training incentives of the employing firm and workers' life-cycle mobility patterns must be taken into account for specific skill accumulation.[3] Furthermore, since skill acquisition increases a worker's

[3]Further assumptions on the model lead to explicit expressions for the functions $g(\cdot)$ and $f(\cdot)$. Assume, for

marginal productivity, it is simple to identify a sufficient condition securing that earnings grow with experience: as long as a worker decides to acquire general skills and her returns to general capital are greater than the period t costs, her earnings should increase as she accumulates experience in the labor force. Similarly, if the worker decides to invest in specific skills and she remains employed at the training firm, her earnings should grow as she accumulates tenure in the same firm (and returns are larger than costs).

Unlike schooling decisions, any post-schooling investment decisions need to consider the employing firm's incentives. In fact, training decisions are best understood as joint considerations between the employing firm and the worker. For general human capital acquisition, Becker (1964) first showed that in labor markets with incomplete contracts, firms would not provide their workers with general skills, as such human capital increases their outside wage offers and credit-constrained workers could not directly bear the cost of general training. The provision of specific capital may be constrained by similar problems: firms lose their investment on trained workers who subsequently leave the firm, whereas workers who finance their training and are promised higher future wages by the firm could be held up, since the acquisition of specific capital is not verifiable by outside parties. As a result, both types of training should be significantly under-provided in flexible labor markets. Nevertheless, Acemoglu and Pischke (1999) found that firms train a large number of apprentices in Germany. To explain this fact, their theoretical framework postulates that in frictional labor markets (e.g., with high unionization rates or large minimum wages), workers could acquire general skills as their outside wage offers are lower than their earnings at the current firm. A similar logic applies to the provision of specific training: in a context in which a worker has the ability to bargain with her employer over earnings, the firm would be unable to renege on the promise of increased earnings for workers who have acquired specific skills. The theoretical arguments presented above indicate that the returns to experience and tenure will depend directly on the degree of labor market

instance, a linearly declining rate of general skill investment with respect to experience ($k_j = g(1 - \frac{\text{Exp}_j}{T})$) and similarly for specific skill investment with respect to tenure ($d_j = s(1 - \frac{\text{Ten}_j}{T})$), where T represents the length of the working life. This structure can be justified from the fact that skill investment should take place early in the career, since there is a longer time period in which workers can accrue the returns to their investment (Ben-Porath, 1967), making it crucially important to understand under which conditions young workers decide to acquire general and/or specific skills in the labor force. Under these assumptions, the relationship between observed earnings, experience and tenure is given by:

$$\ln Y_t = [\ln P_0 - g - s] + \left(\rho_0 g + \frac{\rho_0 g}{2T} + \frac{g}{T}\right) \text{Exp}_t - \frac{\rho_0 g}{2T} \text{Exp}_t^2 + \left(\eta_0 s + \frac{\eta_0 s}{2T} + \frac{s}{T}\right) \text{Ten}_t - \frac{\eta_0 s}{2T} \text{Ten}_t^2,$$

which is an extension of the Mincer model (Mincer, 1974).

frictions present in an economy, which in our context implies that given the institutional features of the labor markets in Brazil an Chile, we should expect higher returns to both experience and tenure in Brazil relative to Chile.

Following our conceptual framework and the existing literature (Abraham and Farber, 1987; Altonji and Shakotko, 1987; Topel, 1991), we estimate equation (2) using polynomials in experience and tenure. These capture potential non-linearities in the accumulation of general and specific human capital. However, estimating equation (2) through ordinary least squares would yield biased estimates of the returns to experience and tenure, as, for instance, higher ability workers tend to spend less time unemployed and are more likely remain employed at better firms for a longer period of time. Abraham and Farber (1987) and Altonji and Shakotko (1987) sought to deal with this problem by pursuing instrumental variable strategies, but, as noted by Buchinsky *et al.* (2010), these strategies fail to fully correct for endogenous mobility, thereby yielding biased estimates of the returns to experience and tenure. As a result, an alternative approach pursued by Topel (1991) proposed estimating a lower bound of returns to experience using a sample of workers starting new jobs, for whom tenure would equal zero. Nonetheless, as this sample includes a large share of voluntary movers, the returns to experience remain biased since this sample of workers changes firms to improve on their previous wages.

An alternative strategy introduced by Dustmann and Meghir (2005) proposed estimating the returns to experience for workers who had been involuntarily displaced from their previous jobs, as these workers would have different levels of pre-displacement experience. However, an OLS estimation of equation (2) using this sample could still yield biased returns to experience and tenure, since individuals with higher returns to experience could have spent more time employed given their potential experience, and similarly, those with higher returns to tenure could have remained with the same firm for longer. As a result, Dustmann and Meghir (2005) proposed correcting for the remaining ability bias within the displaced sample by implementing a control function approach. Their approach is relevant to our research question since the presence of labor market frictions may affect the capacity of high-ability individuals to accumulate experience in the labor market due to the difficulty of finding formal sector jobs and search and matching frictions could affect the capacity of high-ability workers to match and remain with high-paying firms. To estimate the bias in the estimated returns created by labor market institutions, we could use an additional instrument

indicating the degree of labor market frictions in an economy. Nonetheless, this is not a feasible approach in our context since we have data on workers in two different countries and pooling these workers in the same regression would mask country-specific characteristics affecting the returns to human capital. As a result, we generalize equation (2) to allow for individual-specific returns and for country-specific institutions to affect the returns to experience and tenure as follows:

$$ln\ Y_{ict} = \beta_0 + \beta_1^c g(\text{Exp}_{ict}) + \beta_2^c f(\text{Ten}_{ict}) + e_{it}^c g(\text{Exp}_{ict}) + v_{it}^c f(\text{Ten}_{ict}) + \varepsilon_{ict} \tag{3}$$

where c resents a worker's country of employment, and the estimated returns to experience and tenure include a country-specific component, which also captures the relationship between labor market institutions and the rewards to human capital accumulation. We estimate equation (3) for the displaced worker sample separately for each country using both an OLS and a control function approach and compare the estimated returns in the two approaches to capture the degree to which labor market frictions affect high-ability workers' capacity to acquire additional experience and tenure. In our context, we expect that the difference in the estimated returns to experience and tenure under OLS and control function would be larger in Brazil, given the magnitude of the labor market frictions present in the country.

3.2 Empirical Implementation

We estimate equation (3) in two steps. We first identify the returns to experience using the first post-displacement job observation during which tenure equals zero. To correct for the remaining ability bias among displaced workers, we follow a control function approach, using age as an instrumental variable, which provides exogenous variation in experience, as older workers will have accumulated more experience prior to displacement. In this context, the exclusion restriction requires that age does not affect wage offers for individuals with the same (actual and potential) experience, an assumption which is likely to hold in our context as we focus on a narrow age range (20-29 year olds) for workers who have been involuntarily displaced from their job.[4] Moreover, since workers who subsequently re-enter the labor force are those who received acceptable job offers, we correct for endogenous post-displacement labor force participation by introducing an additional

[4]Note that the exclusion restriction is less likely to hold in the general population as older workers may have accumulated additional job search capital, allowing them to receive higher wage offers.

first stage equation, using age as an instrumental variable, as older workers may have accumulated additional job search capital, making them more likely to participate. In our implementation, we interact age variables with potential experience, to account for the fact that workers who have been active in the labor force for longer have had an additional opportunity to accumulate experience. In the participation equation, we also include displaced workers who do not subsequently re-enter the labor force, thereby correcting for extensive margin participation.[5] To identify the model, we further assume that workers cannot predict mass layoffs or closures prior to joining the firm and firms and workers have full information about match quality.

In this context, our first stage for experience (Exp_{it}) and participation (Part_{it}) is specified as follows:[6]

$$\text{Exp}_{it} = \alpha_0 + \alpha_1 age_{it} + \alpha_2 c_{it} + \alpha_3 age_{it} \times c_{it} + \alpha_4 ad_{it} + \alpha_5 (ad_{it} \times c_{it}) + \alpha_6 x_{it} + v_{it}^G \tag{4}$$

$$\text{Part}_{it} = \gamma_0 + \gamma_1 age_{it} + \gamma_2 c_{it} + \gamma_3 age_{it} \times c_{it} + \gamma_4 ad_{it} + \gamma_5 (ad_{it} \times c_{it}) + \gamma_6 x_{it} + v_{it}^P \tag{5}$$

In the two equations above, *age* is a linear term of age, *ad* terms represent dummy variables of age, c_{it} is potential experience and x_{it} is a set of observable variables which includes gender and years of education. The structure of the two first stages implies that conditional on potential experience, older workers have different participation decisions. In fact, since experience equals the sum of past participation decisions, the non-linear specification of age and potential experience allows us to identify both first stage equations. More formally, a rank condition is required, where the vector of the coefficients of excluded instruments in both first stages must have rank two. In our empirical results, we show the rank condition is met in both countries. Following the control function approach proposed by Heckman and Robb (1985), we estimate the first-stage residuals $\hat{v}_{it}^G$, $\hat{v}_{it}^P$ and subsequently interact them with observed experience (Exp_{it}), which allows us to estimate heterogeneous returns to experience in the wage equation, as in our conceptual framework. Lastly, since workers with greater potential experience have been able to acquire more experience, we also interact the residual terms with potential experience (c_{it}). In our control function specifications,

[5] As shown below, this captures fewer than 10 percent of workers do not subsequently re-enter the formal sector in either country.

[6] For notational simplicity, we do not include the country c sub-index, but we estimate the procedure separately for Brazil and Chille.

the residual terms enter linearly and the wage equation is thus estimated as follows:[7]

$$\begin{aligned}\ln w_{it} &= g(Exp_{it}) + x'_{it}\gamma + \delta_1^G \hat{v}_{it}^G + \delta_1^P \hat{v}_{it}^P + \delta_2^G \hat{v}_{it}^G Exp_{it} + \delta_2^P \hat{v}_{it}^P Exp_{it} \\ &+\delta_3^G \hat{v}_{it}^G c_{it} + \delta_3^P \hat{v}_{it}^P c_{it} + \delta_4^G \hat{v}_{it}^G Exp_{it} c_{it} + \delta_4^P \hat{v}_{it}^P Exp_{it} c_{it} + e_{it} \end{aligned} \tag{6}$$

Equation (6) yields unbiased estimates of the returns to experience, and we estimate the returns to tenure in our second step, first adjusting observed wages by subtracting the growth in wages explained by experience:

$$\widetilde{\ln w_{it}} = \ln w_{it} - \widehat{g(Exp_{it})} \tag{7}$$

where $\widehat{g(Exp_{it})}$ is estimated from equation (6). While residual wages could then be regressed against a polynomial of tenure, $f(Ten_{it})$, post-displacement tenure remains an endogenous variable, as workers decide whether to remain in the firm or switch in every period. As a result, we estimate a reduced form for tenure, in which we also use age as an instrumental variable, following from the fact that age may affect mobility decisions for workers on the basis of family and other life-cycle considerations. In fact, the tenure reduced form is estimated in the same manner as those for experience and participation, allowing for non-linearities in age and an interaction with potential experience.[8] [9] As in Dustmann and Meghir (2005), we estimate the returns to tenure using the first post-displacement job spell of displaced workers and regress wages on a polynomial of tenure and the all residuals from the experience, participation and tenure first stages, interacted by experience, tenure, and sector tenure (for notational simplicity, v_{it}^{RES} includes the full set of eighteen residual terms). The estimating equation is therefore:

$$\widetilde{\ln w_{it}} = \ln w_{it} + f(Ten_{it}) + v_{it}^{RES} + e_{it} \tag{8}$$

[7] While Dustmann and Meghir (2005) estimate their model including the returns to sector tenure, we have not included sector tenure in our specification given our interest in the importance of general- and specific- capital. We have separately estimated the model allowing for sector tenure and the estimated returns to experience and tenure in Brazil and Chile do not change significantly. These results are available upon request.

[8] Unlike Dustmann and Meghir (2005), we estimate each first stage by restricting the estimation to the sample of workers who are included in the subsequent wage equation. We impose this restriction since the exclusion restrictions discussed above are more likely to hold in each of these samples. As a result, the experience and participation first stages are estimated using displaced workers' first post-displacement job observation, and the tenure first stage includes these workers first job spell.

[9] The additional first stage equation for tenure implies that the coefficients on the excluded instruments in the three reduced form experience, participation and tenure equations must have rank three.

The estimated returns to tenure follow from the empirical estimates of equations (6) and (8) in both Chile and Brazil. As suggested by the authors, we estimate our standard errors after bootstrapping the entire estimation procedure. We compare the results from the control function approach to different least squares specifications to carefully show how failing to account for the endogeneity of experience and tenure results in biased estimates. Moreover, we implement formal tests of the difference in the estimated returns to experience and tenure from the least squares and control function specifications in each country to characterize the interaction between labor market frictions and the rewards to human capital accumulation. We next introduce the data sources used in Brazil and Chile and discuss the construction of the variables used in the empirical analysis.

4 Data Sources and Summary Statistics

4.1 Data Sources

Chile In Chile, the Unemployment Insurance (UI, *Seguro de Cesanta*) database contains matched employee-employer data for all formal sector employment contracts. The UI database has records of all formal workers' monthly earnings from November 2002 through 2013, including upwards of seven million workers. UI provides information on workers' monthly earnings, sector of employment, state of residence, and observable characteristics, such as educational attainment, gender and age. For every job held by every worker, we observe the month of entry and exit, which we use to construct a measure of months worked in each year. UI data therefore allows us to construct a panel of all formal sector workers in Chile during 2003 through 2012. Furthermore, as UI includes a unique establishment-level identifier, we also construct a longitudinal panel of the universe of firms in Chile. Upon combining the worker and firm panels, we are able to track worker flows across establishments, allowing us to directly measure experience and tenure.

Brazil In Brazil, the Relacao Anual de Informaes Sociais (RAIS) database contains matched employee-employer data from a mandatory annual survey filled by all registered firms in the formal sector. In this paper, we use RAIS data for the country's 27 states for 2003 through 2012.[10] As the Ministry of Labor has been known to levy fines on inaccurate reports, firms tend to hire specialized accountants to ensure the correct completion of the RAIS survey, resulting in highly accurate data

[10] By 2002, the survey covered more than 95 percent of the universe of formal sector businesses and workers.

(Moser and Engbom, 2016). RAIS includes unique, time-invariant person identifiers, which allows us to construct a panel of formal sector workers in the period of interest. Moreover, the dataset includes individual-level characteristics such as age, gender, educational attainment, nationality and state of residence.

For each job held by every formal sector worker, we observe the number of total days employed in each year, including the month of entry and exit (aggregated to construct a variable of total months worked). The dataset includes earnings information on the average monthly compensation received by each worker, which includes regular salary payments, holiday bonuses, performance-based and commission bonuses, tips, and profit-sharing agreements. Throughout the analysis, we focus on a worker's primary employer, defined as the one in which the worker earns the highest total earnings in a year. RAIS also includes a unique establishment-level identifier, which allows us to construct a longitudinal panel representing the universe of establishments and firms in Brazil over the 2003-2012 period. Lastly, we combine the employee and employer panels to track worker flows across establishments and firms to carry out our empirical strategy.[11]

Variable Construction To develop comparable insights across the two countries, we restrict our analysis to workers born in 1983-84, for whom we observe their labor market trajectories from entry through age thirty.[12] The sample size in Chile includes 241,207 workers, the majority of whom enter the formal sector in 2003 (42.2 percent) or in 2004 (20.2 percent). In Brazil, the initial sample size included 408,283 workers, but to keep the sample sizes comparable, we carry out our analysis using a random sample of 250,000 workers. We observe similar patterns with respect to the year of entry: 54.3 percent of the sample enters in 2003 and 15.1 do so in 2004. Our administrative data sources allow us to construct uniform measures of potential experience, experience, tenure, and labor market participation across the two countries. Following the literature, we define potential experience, c_{it}, as age at time t minus education minus six. Experience is defined as the number of months (defined as year-equivalents) that worker i has worked prior to starting period t. Tenure is similarly defined as the number of months worker i has worked in firm f up through year t. Meanwhile, the participation variable is defined as the number of months worked in the formal

[11] We exclude observations which are missing either the firm or the individual-level identifier or those with missing values for earnings or dates of employment.

[12] Given the large wage premium enjoyed by public sector workers in Latin America, we exclude workers who transition through this sector from our analysis.

sector in year t, in year-equivalents. Our wage measure reflects monthly earnings at the primary employer. This measure is first expressed in real terms (deflated by each country's Consumer Price Index), and subsequently converted into US dollars for clearer comparability across the two countries.

The two administrative data sources allow us to identify the month during which a worker switches from one firm to another, exits, or re-enters the formal labor market. To correctly capture workers who have been involuntarily displaced, we focus on distressed firms which close down as well as those laying off a large share of their workers. Firm closures are events in which a firm with at least 50 employees at its peak closes down and whose identifier does not subsequently re-appear in the country's administrative database. Similarly, mass layoffs include events in which an establishment's employment levels drop below 50 percent relative to its pre-layoff peak, and whose employment does not subsequently recover. We impose the 50 percent threshold, which is more stringent than the developed country literature standard, to avoid including employers which selectively lay off workers or those with a large share of voluntary movers. Furthermore, the 50-employee requirement avoids capturing establishments which undergo a small change in employment levels, while showing a large percentage drop in employment.[13] Our sample includes displaced workers whose initial firms downsized in either the same year or in the following one. In Chile, 33,008 workers in our sample were displaced in a mass layoff event and 11,003 were let go when their firm closed down, and 97.7 percent of workers in the first group and 97.5 percent of workers whose firms closed down eventually re-entered the formal sector. Meanwhile, in Brazil, 22,897 individuals were affected by a mass layoff and 9,029 by a firm closure event. 87.7 percent of workers in the first group re-entered the labor force and 87.5 percent in the latter did so as well. While the re-entry rates in Brazil are lower than in Chile, they significantly exceed the 60 percent re-entry rate for all displaced workers in Brazil (Saltiel, 2017), likely reflecting the frictional nature of the Brazilian labor market. As mentioned above, we estimate the returns to tenure from the sample of displaced workers who eventually re-entered the formal sector, yielding a sample of 43,355 workers in Chile and 27,975 individuals in Brazil.

Lastly, while our data sources allow us to cover the universe of formal sector workers, we are

[13] We have conducted robustness tests on the choice of the firm size and percent change thresholds and found the main results are robust across specifications.

unable to observe workers' participation in the informal sector, an important consideration in any developing country. While this sector is not particularly large in Chile, the rate of informality in Brazil reached 49 percent in 2010, among highest in Latin America (Dix Carneiro and Kovak, 2016). Using longitudinal data in Chile, (López García, 2015) has previously found that experience in the informal sector does not result in wage gains in formal jobs. Given our data availability, we impose this assumption in our empirical analysis, but we acknowledge that further evidence is needed about the transferability and returns to human capital across these two sectors.

4.2 Summary Statistics

Given the lack of evidence about young workers' labor market trajectories in Latin America, we first present various descriptive statistics for both Brazil and Chile in Table 1. As Chile's GDP per capita far exceeds that of Brazil, it is unsurprising that both initial and final salaries are higher for young workers in Chile relative to those in Brazil. For instance, young Chilean workers employed in the formal sector in 2003 earned U$D 380 on average, while Brazilian workers earned an average of U$D 175 per month. At the same time, in Figure 1, we show that both groups of workers experienced significant growth by 2012, resulting in monthly earnings more than doubling over this time period. In terms of other observable characteristics in Table 1, the sample in both countries skews towards males, and on average, workers in both countries had attained less than a high school degree, also representative of educational attainment in these two countries.

Table 1 also presents preliminary evidence on young workers' mobility. We find that while workers in both countries accumulate similar rates of formal sector experience by 2012 (3.75 years in Brazil and 3.62 years in Chile), Chilean workers are more mobile, holding an average of 4.00 jobs through 2012, with Brazilian workers having 2.97 different jobs. These numbers seem low relative to the U.S. context, where Topel and Ward (1992) found that during the first ten years of labor market participation, a young worker holds an average of seven jobs and Neal (1999) found similar patterns using NLSY79 data. On the other hand, the measure of number of jobs held presented above excludes young workers' participation in the informal sector, resulting in a likely underestimate of the number of total jobs held by workers in our sample. Figure 2 confirms the fact that young workers in Chile sample more jobs than Brazilian workers, regardless of the year of entry into the formal sector, fitting in with the evidence of more flexible labor market regulations

in Chile.

Figure 3 shows annual wage growth by workers' age in both countries. The wage growth measure is constructed by averaging the the growth rate of real monthly earnings for all individuals observed in adjacent years. We find that in both countries, wage growth rates are fairly stable across the first decade of formal sector participation, ranging between 6 and 11 percent. Note that the time period of interest reflected strong economic growth in both countries, which may explain why young workers' wages continued growing at high rates as they approached thirty years old. Lastly, in Figures 4 and 5, we decompose wage growth by workers' experience and mobility status. In both Brazil and Chile, we find that large wage gains for workers in their first year of labor market experience, with decreasing (yet positive) gains for more experienced workers, reflecting lifecycle wage growth considerations. In terms of mobility, we find that Brazilian workers who stay at the same firm enjoy slightly larger gains relative to movers, but the difference is not statistically significant for more experienced workers. In Chile, on the other hand, stayers' average wage growth far outpaces that of movers, possibly reflecting greater returns to tenure, or differential mobility by ability levels relative to Brazil. As these figures do not allow us to properly estimate whether experience or tenure is the driving force behind wage growth, we estimate our control function strategy in the next section and present our main results.

5 Estimated Returns to Experience and Tenure

Returns to Experience. Table 2 presents the cumulative returns to experience in the first five years in the labor force for workers in Brazil and Chile. We present the results from our control function approach along with three least squares estimates, where the first specification estimates the returns using the full sample of workers (OLS All), the second one follows Topel (1991) and estimates the return to experience in sample of workers starting a new job and the return to tenure based on within firm growth (OLS NJ), and the third one includes the same sample restrictions as in the control function approach, but does not directly control for endogeneity of experience and mobility decisions (OLS PD).

This method allows to correct for two different types of endogeneity. On one hand, we have endogeneity coming from different types of workers selecting themselves to certain jobs and accu-

mulating experience and tenure (individual bias). On the other hand, as our conceptual framework explains, workers may accumulate experience or tenure differently in presence of more frictions (institutional bias). The comparison between the estimations with the full sample, and the estimations with the sample of post-displaced workers (columns (1) and (3)) are an indication of the direction of the individual bias, while the comparison between the third and the fourth column allows us to discuss the direction of the institutional bias.

Panel A presents the returns to experience in Brazil. Across all four specifications, we find increasing returns to experience throughout first five years in the labor force.[14] The full sample OLS estimate shows returns of 5 percent in the first year of experience, eventually reaching 24 percent by the fifth year in the labor force. Upon restricting the sample to workers in new jobs, we find a smaller return to experience through five years, in the range of 20 percent. On the other hand, when we limit our sample to workers who have been involuntarily displaced, we find larger returns in both the first year and through the fifth year, eventually reaching 31 percent. The last column presents the estimates from our preferred control function approach. In Appendix Table A, we present the coefficients from the first stage regressions in both countries. In Brazil, the p-values for the three first stages are statistically significant at the 0.001 level (a necessary condition for the rank condition), which implies that correcting for endogenous experience and tenure is an important consideration in our context.[15] As a result, relative to the estimates in column (3), our control function specification adds the estimated residuals from the experience and participation reduced forms, resulting in significantly higher returns in the first year, reaching upwards of 10 percent, and larger cumulative returns through year five. As a result, we find that accumulating five years of experience in the formal sector in Brazil generates an additional wage gain in the range of 43 percent. The estimated marginal returns to experience in Brazil have not decreased even after a worker has completed five years in the labor force, which is in line with the descriptive statistics presented in Figure 4, where we showed significant increases in earnings for workers with more than four years of experience.

[14] Our control function approach estimates the returns to experience using the first post-displacement job observation of young workers who have been involuntarily fired from their job. As a result, the sample used in columns (3) and (4) is relatively unexperienced, such that fewer than 1% of them have more than 5 years of experience. As a result, we only present estimated returns to year 5 to avoid making out-of-sample predictions.

[15] Dustmann and Meghir (2005) note that a formal test of the rank condition requires that the rank of the excluded instruments in the three first stages must have a maximum value of three. We follow their approach and implement a test with a null hypothesis of rank two and reject it against the required rank three.

Panel B presents the estimated returns to experience in Chile. The first two OLS specifications show large returns to the first year of experience and the returns remain elevated through the fifth year, exceeding upwards of 35 percent. However, when we restrict our sample to displaced workers, the estimated returns in the first five years become significantly larger. On the other hand, including the residuals from the two first stage regressions does not deliver significantly different returns. The coefficients on the first stage residuals are presented in Appendix Table B, which show that older workers have accumulated more experience in the labor force, and that workers are most likely to participate in the formal sector in their mid-20s. Lastly, we note that the estimated returns to experience in Chile are larger than the returns in Brazil in the first five years in the labor force.

In our conceptual framework, we had argued that returns to experience could be larger in a frictional labor market like Brazil under certain assumptions. Nonetheless, the empirical evidence presented in Column 4 of Table 2 shows that by the fifth year of formal sector experience, Chilean workers experience larger wage gains than their Brazilian counterparts. While the interpretation of these results depends on the crucial assumption of equivalent secular returns to skill across both countries, our empirical results seem to indicate that in a more flexible labor market, in which young workers sample more jobs, there are larger returns to general human capital. Furthermore, the comparison between columns shows the direction of both the individual bias and the institutional bias is downward in both countries, since returns increase as we estimate with the sample of post-displaced workers, and as we implement the control function approach. As expected, the magnitude of the difference between columns (3) and (4) is larger in Brazil than in Chile, because in presence of labor market frictions is stronger in Brazil. However, performing a Wu-Hausman test on the differences between these two columns for both countries shows that there are no statistical differences in either country. This is expected since experience is more dependent on the worker's job history than on the way they are able to find jobs after being displaced in a certain institutional framework.

We note that despite focusing on a different context than the existing literature, our estimated returns to experience are similar to those presented by Dustmann and Meghir (2005). The authors' estimates of the returns to experience for unskilled and skilled young workers in Germany show that the first year of experience delivers wage gains of 9% and 7%, respectively, but the marginal returns

fall to around zero by the fifth year. The rest of the literature estimates the returns using survey data and including workers of all ages, making comparability a potential issue. Nevertheless, our estimated results remain in the range of the previous literature. For instance, Dustmann and Pereira (2008) find that British and German workers enjoy wage gains of 45 and 20 percent, respectively, following their first five years in the labor force and Topel (1991) finds that accumulating five years in the U.S. labor force results in wage gains in the range of 20 percent. We note, however, that the returns may be larger in the two countries in our sample vis-a-vis their developed country counterparts due to certain labor market idiosyncrasies affecting young workers' trajectories, such as the structure of government-sponsored training programs in Chile and the role of apprenticeships in Brazil (Rodriguez *et al.*, 2017; Corseuil *et al.*, 2013).

Returns to Tenure. In Table 3, we present the estimated cumulative returns to the first five years of tenure in Brazil and Chile from the four specifications discussed above. Panel A presents the results in Brazil, where across all specifications, we find small, yet positive and statistically significant returns to tenure. We find similar returns across our three OLS specifications, where the estimated return to staying with the same firm for one year is in the range of 1-1.5 percent, eventually climbing to 5-8 percent for workers who reach five years of tenure. Recall the estimated returns in column (2) follow from an estimation based on within-firm growth from all jobs, such that workers employed for one period are dropped from the sample. The fact that the returns are similar across the two columns can be explained by the fact that workers who leave their jobs include a mix of voluntary and involuntary movers. In fact, upon restricting our sample to involuntarily displaced workers, the estimated annual returns to tenure fall to about one percent. Our control function estimate, which controls for the endogeneity in post-displacement mobility decision, yields significantly larger returns to tenure. As a result, the first year of tenure is associated with an earnings increase of 4.2 percent, reaching 14.8 percent after five years in the same firm. The significant increase in the estimated returns from column (3) to (4) implies that failing to account for the endogeneity in post-displacement mobility decisions results in an downward-biased estimate of the returns to tenure, similar to the findings in (Dustmann and Meghir, 2005).

Panel B presents the returns to tenure in Chile. Across the full sample OLS specification, the initial returns to tenure are higher than in Brazil, in the range of 4 percent in the first year with the firm, eventually climbing to upwards of 20 percent percent for workers who reach five years of

tenure. The estimated coefficients fall significantly as we move from the full sample estimate in column to the second and third columns, which implies that in Chile, those who stay at the firm likely had better relative outcomes to workers who switched employers. Lastly, in our preferred control function approach, we find slightly larger returns of 2.1 percent in the first year, reaching 10 percent after spending five years with the same firm. As in Brazil, the estimated returns to tenure increase once we control for the endogeneity in post-displacement mobility, a result which is confirmed by the coefficients on the tenure residuals from the first-stage, presented in Appendix Table B.

Comparing columns (3) and (4) in this case shows that the institutional bias goes in the downward direction. Again,the magnitude of the difference between columns (3) and (4) is larger in Brazil than in Chile, due to the presence of greater labor market frictions. The Wu-Hasuman test on the differences between these two columns for both countries show that there are statistical differences in both countries. As expected, in Brazil differences are greater than in Chile. The difference in returns to one through five years of tenure is greater in both countries up to a confidence level of 99%. Since tenure is obtained only after workers have been displaced, the institutional bias is more visible in this case.

Our estimated returns to tenure in both Brazil and Chile fall well within existing estimates in developed countries. Dustmann and Meghir (2005) find that the first year of tenure is associated with wage gains of 2.4 percent for unskilled workers and of 4 percent for skilled workers in Germany, eventually falling to 1.7 percent and 1.1 percent after five years, respectively. Meanwhile, Dustmann and Pereira (2008) find smaller returns to tenure, as cumulative returns reach 4 percent in the United Kingdom and 0-7 percent in Germany after five years with the same firm. Lastly, the estimated returns in the U.S. context are also in line with those presented in this paper, given Altonji and Williams' (2005) estimated return of 11 percent after ten years of tenure and Topel's (1991) estimate of 28 percent for ten years of job seniority. In our conceptual framework, we noted the returns to tenure could be larger in a more regulated labor market like Brazil, under the assumption of equal returns to specific capital across both countries. Again, while our interpretation depends on the validity of this assumption, we find preliminary evidence than in a country with stricter labor market regulations, where workers sample fewer jobs (as shown in Figure 2), there are larger returns to remaining at the same firm. Given our theoretical set-up, this result could be explained

by the fact that both firms and workers have an incentive to accumulate specific skills given the larger frictions to young workers' mobility in Brazil.

6 Results by Educational Attainment

To better understand how these two countries reward experience and tenure, we further explore the heterogeneous returns by educational attainment in Brazil and Chile. Given the recent boom in post-secondary enrollment in Latin America, focusing on the returns by levels of educational attainment is of particular policy interest, as the enrollment rate has surged from 17 percent in 1991 to 43 percent by 2012 (Espinoza and Urzúa, 2016). This trend has also taken place in both countries in our sample. In Brazil, 60 percent of 25-35 year olds had not completed a high school degree in 1995, but this share has since fallen exponentially, down to 36% by 2008. Moreover, the share of workers in this age range with a tertiary education has almost doubled over the same time period, from 8 percent in 1995 to 14 percent in 2008. Despite starting from higher levels of educational attainment, Chile has experienced a similar trend in the past two decades. The share of 25-35 year olds with less than a high school degree fell from 20 percent in 1994 to 10 percent by 2006, while the share of tertiary education finishers increased from 16 percent to 27 percent over the same time period (Aedo and Walker, 2012). At the same time, there are large returns to an additional year of education in both countries, in the range of 10 percent in Brazil and 12 percent in Chile. Furthermore, the returns to completing a five-year university degree yield an estimated wage gain of 92 percent in Brazil relative to just graduating from high school and of 112 percent in Chile. Nevertheless, this recent educational expansion has possibly resulted in declining quality of institutions and students. As a result, the estimated return to education has fallen in Brazil, from 13 percent in 1995 to 10 percent in 2008, while remaining flat in Chile. In this context, understanding whether increased labor market experience or increased tenure could deliver greater wage gains for young workers is a particularly policy relevant question.

Table 4 presents the estimated returns to experience and tenure from our control function approach, using the first post-displacement job observation for experience and the full post-displacement spell for tenure. Since the two administrative datasets report workers' final educational attainment, we directly group them into one of three relevant categories: workers with less than a high school

degree, high school completers, and those who have gone beyond high school. Since there are different patterns of educational attainment across the two countries, we note that a direct comparison of the returns for the same educational group across the two countries may not be the relevant parameter of interest. Nonetheless, comparing the returns across educational groups in each country provides guidance about the relative returns to experience and tenure.

Panel A presents the estimated returns to experience. In Brazil, we find that the first year of formal sector experience results in wage gains in the range of 8-12 percent for workers in each of the three educational groups. After five years in the labor force, the returns remain large, in the range of 35-50 percent, and do not appear to have yet flattened out. We find larger returns to experience for more educated workers, and the difference in returns between the two less educated groups is not statistically significant. In Chile, we find larger returns to initial experience across the three groups, in the range of 15 percent, eventually reaching 50-60 percent after five years in the labor force. Similar to the Brazilian case, we find that workers who have not gone beyond high school enjoy smaller returns to formal sector experience, whereas the returns to the most educated group exceed 60 percent after five years. We note that despite the different institutional setting, the heterogeneous returns to experience by educational categories in Brazil and Chile resemble those found by Dustmann and Meghir (2005) for young workers in Germany, where there are larger returns to experience for unskilled workers vis-a-vis their skilled counterparts.

In Panel B, we show the heterogeneous returns to tenure by educational attainment. In line with the returns to experience, we find different returns across the three groups in both countries. In Brazil, the first year returns to remaining at the same firm are in the range of 4-5 percent across the three groups, but by the fifth year they have climbed to 20 percent for workers who have not gone beyond high school and upwards of 25 percent for the highest educated workers. In Chile, we find similar returns. The returns to tenure for the three groups are statistically significant in the first five years, but there are differences in the magnitudes as early as the first year: for workers in the first two groups, initial returns are in the range of 2 percent, but exceed 6 percent for workers who have gone beyond high school. This difference persists through the fifth year, resulting in cumulative returns of 11 percent for workers with less than a high school degree, whereas the returns for those who have gone beyond high school exceed 35 percent. Within our conceptual framework, the finding that the returns to both experience and tenure are larger for the highest educated workers

in both countries implies that firms in Brazil and Chile provide their highest educated workers with more specific as well as general training, resulting in higher returns to experience and tenure. Nonetheless, the recent educational expansion that has taken place in both Brazil and Chile may imply that the new cohort of college-educated workers may enjoy different returns than those in the 1983-84 cohort. Lastly, we note that as the returns to formal sector experience for less educated workers in both countries, policymakers should design policies based around helping young workers accumulate experience in the formal sector.

7 Conclusion

Since the first years of a workers' labor market experience account for a major share of their lifetime wage growth, gaining a better understanding of the driving forces behind this phenomenon is of crucial importance, especially as governments implement various policies aimed at helping these workers. In this paper, we have provided the first empirical estimates to experience and tenure in Latin America, by taking advantage of matched employee-employer data in two countries, a first in the developing country literature. As the existing literature had largely focused on developed countries, our focus provides an important contribution to this research agenda. In our analysis, have examined both conceptually and empirically how the returns to experience and tenure differ in the face of different regulatory contexts across Brazil and Chile. Furthermore, as our data sources allow us to accurately identify labor market transitions, we have focused our empirical approach on involuntarily displaced workers, and further implemented a control function approach to corrects for endogenous mobility and participation decisions by focusing on involuntarily displaced workers.

Our empirical results show that in a more frictional labor market like Brazil, where young workers sample fewer jobs than in Chile, workers enjoy larger returns to tenure than in Chile, where workers instead enjoy larger returns to experience. We interpret this result within our conceptual framework, which implies that firms in more frictional labor markets provide their workers with more specific training than in a more lax background. While this interpretation is subject to as assumption about the underlying rates of return to skill in both countries, the evidence presented in this paper represents a first approximation towards achieving a better understanding of the factors which contribute to earnings growth for young workers in Latin America. As we have found differ-

ent returns to experience and tenure in Chile and Brazil, and we have further found heterogeneous returns for workers with different levels of educational attainment, we note that one policy may not be adequate to address the needs of young workers in the region, but rather focused approaches should prevail instead.

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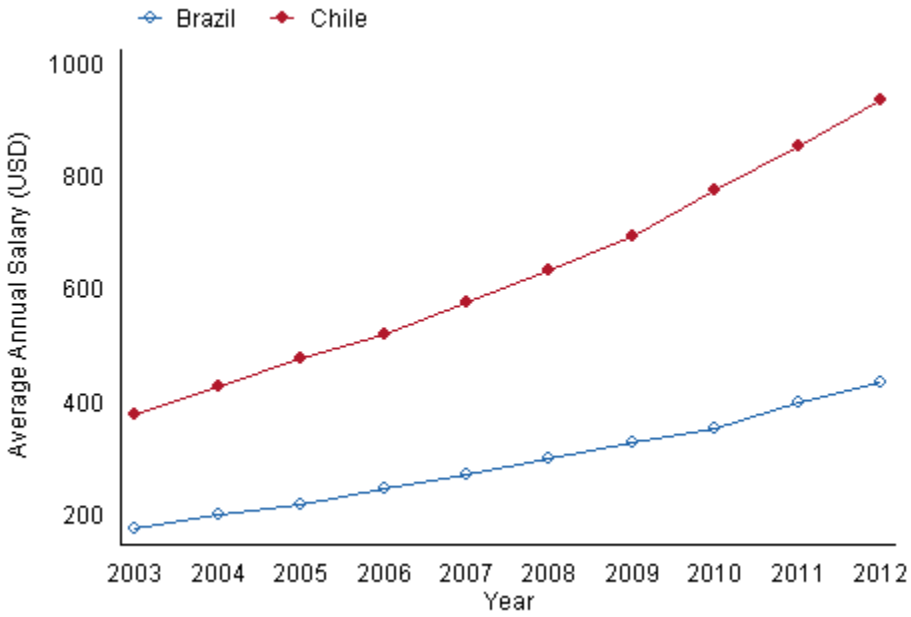
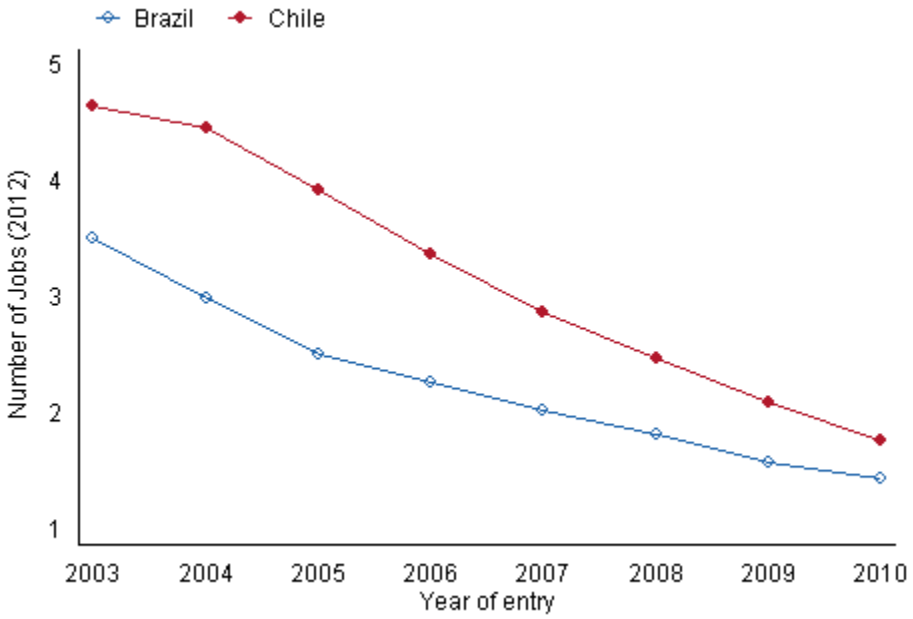
Figure 1: Average Monthly Earnings in Brazil and Chile. 2003-2012



Notes: Figure 1 includes average monthly wages for all workers who are observed in the administrative datasets of each country in each year from 2003-2012. The values represent average monthly earnings in US Dollars, which are calculated using the yearly exchange rate in each country relative to real earnings in Reais (in Brazil) and Chilean Pesos (in Chile).

Sources: RAIS and Chilean Unemployment Insurance Database

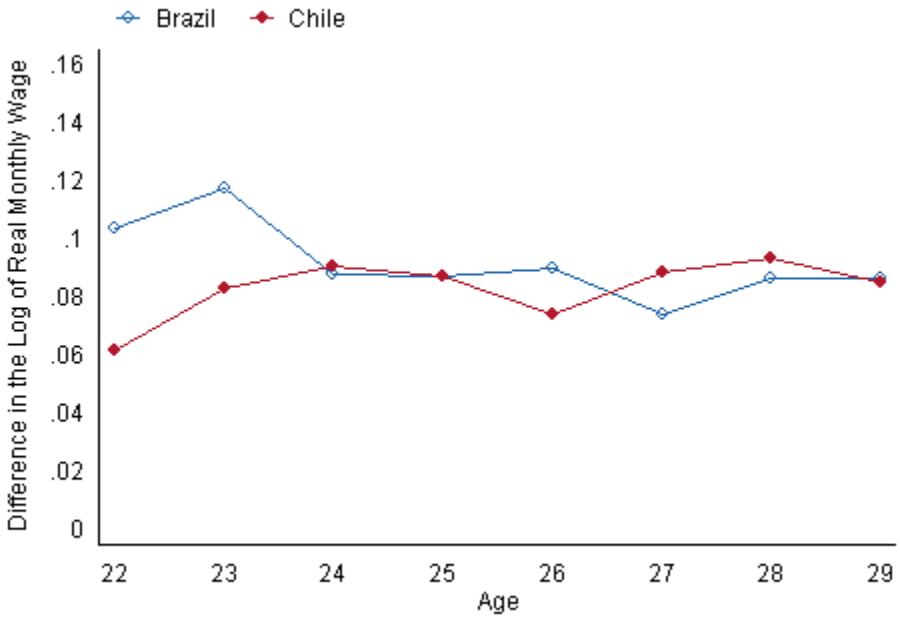
Figure 2: Total Number of Jobs Held Through 2012 by Workers in Brazil and Chile



Notes: Figure 2 shows the total number of jobs held through 2012 for all workers in our sample, who entered the labor force in each year from 2003 through 2010

Sources: RAIS and Chilean Unemployment Insurance Database

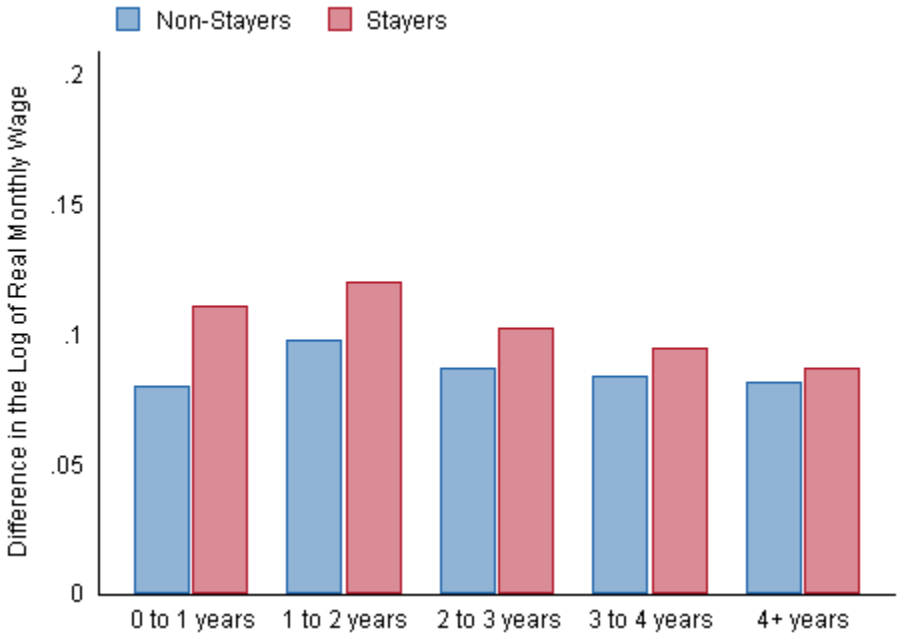
Figure 3: Average Monthly Earnings Growth by Age in Brazil and Chile



Notes: Figure 3 shows the average wage growth in Brazil and Chile for workers whose age is between 22 and 29 in the first year from which such growth is calculated.

Sources: RAIS and Chilean Unemployment Insurance Database

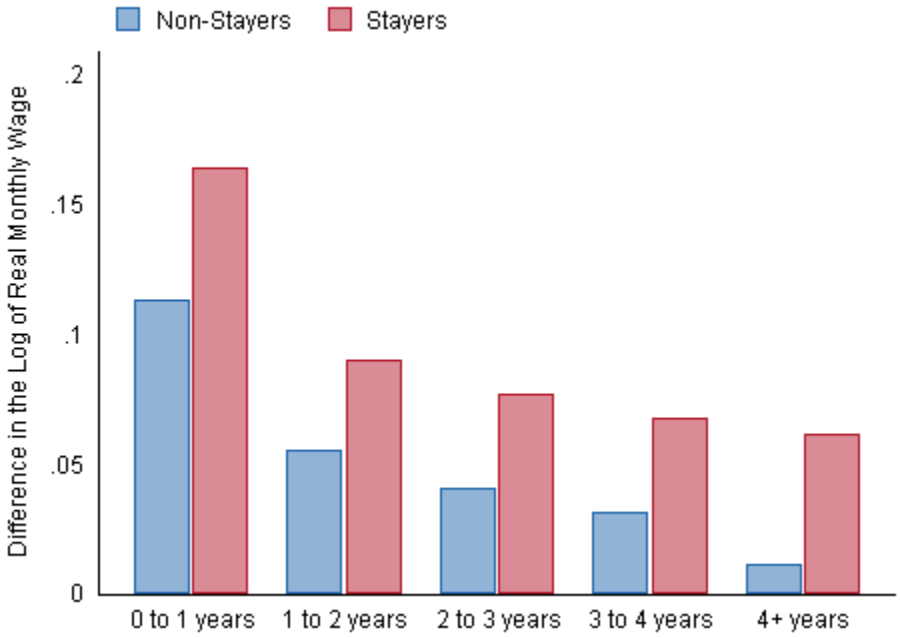
Figure 4: Average Monthly Earnings Growth by Experience and Mobility Patterns in Brazil



Notes: Figure 4 shows the average wage growth in Brazil for workers who have accumulated experience to land in each of the five bins and further divides it by workers' mobility decisions, where stayers include workers who remain with the firm and non-stayers represent workers who work in a different employer than in the previous year.

Source: RAIS

Figure 5: Average Monthly Earnings Growth by Experience and Mobility Patterns in Chile



Notes: Figure 5 shows the average wage growth in Chile for workers who have accumulated experience to land in each of the five bins and further divides it by workers' mobility decisions, where stayers include workers who remain with the firm and non-stayers represent workers who work in a different employer than in the previous year.

Source: Chilean Unemployment Insurance Database

Table 1: Summary Statistics

	Brazil	Chile
Monthly Salary in 2003 (USD)	175.25	380.44
	(144.07)	(183.11)
Monthly Salary in 2012 (USD)	435.50	938.98
	(362.90)	(693.70)
Years of Education	11.24	11.61
	(2.82)	(2.82)
Share of Males	0.65	0.67
	(0.48)	(0.47)
Number of Jobs	2.97	4.00
	(1.75)	(1.91)
Total Experience	3.74	3.62
	(2.70)	(2.10)
Observations	243,414	238,157

Note: Standard Deviation in parenthesis. Wages are first deflated by each country's consumer price index (2002 values as base), and then converted into USD using annual exchange rates. The number of jobs held variable reflects worker's primary employer, such that one worker holds at most one job a year.

Sources: RAIS and Chilean Unemployment Insurance Database

Table 2: Returns to Experience

Panel A. Brazil

	OLS All	OLS NJ	OLS PD	CF PD
1 Year	0.052	0.041	0.069	0.102
	(0.001)***	(0.002)***	(0.011)***	(0.023)***
2 Years	0.100	0.079	0.131	0.187
	(0.002)***	(0.003)***	(0.015)***	(0.042)***
3 Years	0.146	0.117	0.189	0.264
	(0.002)***	(0.003)***	(0.016)***	(0.060)***
4 Years	0.192	0.157	0.247	0.341
	(0.002)***	(0.003)***	(0.015)***	(0.079)***
5 Years	0.239	0.203	0.311	0.425
	(0.002)***	(0.003)***	(0.015)***	(0.100)***
Observations	1,526,581	478,936	27,975	27,975

Note: Standard Errors in Parenthesis;
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Panel B. Chile

	OLS All	OLS NJ	OLS PD	CF PD
1 Year	0.123	0.136	0.184	0.163
	(0.001)***	(0.002)***	(0.011)***	(0.022)***
2 Years	0.214	0.239	0.321	0.296
	(0.002)***	(0.003)***	(0.016)***	(0.040)***
3 Years	0.281	0.319	0.424	0.405
	(0.002)***	(0.004)***	(0.017)***	(0.056)***
4 Years	0.330	0.382	0.506	0.495
	(0.002)***	(0.004)***	(0.017)***	(0.075)***
5 Years	0.368	0.438	0.579	0.574
	(0.003)***	(0.004)***	(0.019)***	(0.103)***
Observations	1,707,803	714,565	43,355	43,355

Note: Standard Errors in Parenthesis;
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

The estimated returns to experience represent the cumulative returns a worker experiences after reaching t years of participation in the formal sector. We estimate the results using a polynomial of degree three in experience.
OLS All: Estimated returns follow from the full sample of workers in the sample.
OLS NJ: Estimated returns are derived from first observation at new jobs, excluding initial firm.
OLS PD: Returns to experience are estimated using the first observation in the first post-displacement job.
Control Function: Returns to experience are estimated using the first observation in the first post-displacement job.

Table 3: Returns to Tenure

Panel A. Brazil

	OLS All	OLS NJ	OLS PD	CF PD
1 Year	0.016	0.017	0.010	0.042
	(0.000)***	(0.000)***	(0.001)***	(0.007)***
2 Years	0.033	0.034	0.019	0.072
	(0.001)***	(0.001)***	(0.003)***	(0.007)***
3 Years	0.050	0.052	0.029	0.097
	(0.001)***	(0.001)***	(0.004)***	(0.007)***
4 Years	0.067	0.070	0.039	0.118
	(0.001)***	(0.001)***	(0.006)***	(0.007)***
5 Years	0.084	0.088	0.049	0.148
	(0.002)***	(0.002)***	(0.008)***	(0.007)***
Observations	1,526,581	805,051	57,612	57,612

Note: Standard Errors in Parenthesis;
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Panel B. Chile

	OLS All	OLS NJ	OLS PD	CF PD
1 Year	0.042	0.009	0.012	0.021
	(0.000)***	(0.000)***	(0.002)***	(0.006)***
2 Years	0.085	0.018	0.023	0.038
	(0.001)***	(0.001)***	(0.003)***	(0.006)***
3 Years	0.131	0.027	0.035	0.063
	(0.001)***	(0.001)***	(0.005)***	(0.006)***
4 Years	0.178	0.036	0.047	0.085
	(0.002)***	(0.002)***	(0.007)***	(0.006)***
5 Years	0.227	0.045	0.059	0.100
	(0.003)***	(0.002)***	(0.009)***	(0.007)***
Observations	1,707,803	755,081	79,529	79,529

Note: Standard Errors in Parenthesis;
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

The estimated returns to tenure represent the cumulative returns a worker experiences after reaching t years of tenure at the same firm. We estimate the results using a polynomial of degree three in tenure.
OLS All: Estimated returns follow from the full sample of workers in the sample.
OLS NJ: Returns are calculated from within firm wage growth, as the differences in logs of wages over time.
OLS PD: Returns to tenure are estimated using the full spell of the first post-displacement job.
Control Function: Returns to tenure are estimated using the full spell of the first post-displacement job.

Table 4: Heterogeneous Returns to Tenure and Experience by Education Level

Panel A. Experience

	Brazil			Chile		
	Less than High School	High School Graduate	More than High School	Less than High School	High School Graduate	More than High School
1 Year	0.093	0.086	0.120	0.148	0.160	0.155
	(0.024)***	(0.022)***	(0.026)***	(0.022)***	(0.021)***	(0.025)***
2 Years	0.175	0.166	0.229	0.274	0.296	0.293
	(0.045)***	(0.042)***	(0.050)***	(0.040)***	(0.040)***	(0.047)***
3 Years	0.246	0.242	0.327	0.380	0.408	0.414
	(0.064)***	(0.060)***	(0.070)***	(0.056)***	(0.057)***	(0.065)***
4 Years	0.305	0.312	0.415	0.465	0.496	0.519
	(0.084)***	(0.078)***	(0.088)***	(0.075)***	(0.076)***	(0.085)***
5 Years	0.354	0.376	0.491	0.528	0.560	0.607
	(0.108)**	(0.099)***	(0.107)***	(0.105)***	(0.102)***	(0.113)***
Observations	7,436	16,405	4,134	7,436	16,405	4,134

Note: Standard Errors in Parenthesis;
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Panel B. Tenure

	Brazil			Chile		
	Less than High School	High School Graduate	More than High School	Less than High School	High School Graduate	More than High School
1 Year	0.037	0.037	0.048	0.022	0.029	0.066
	(0.006)***	(0.006)***	(0.007)***	(0.007)**	(0.006)***	(0.008)***
2 Years	0.075	0.076	0.099	0.044	0.059	0.136
	(0.013)***	(0.012)***	(0.014)***	(0.014)**	(0.012)***	(0.016)***
3 Years	0.115	0.115	0.151	0.067	0.090	0.211
	(0.021)***	(0.018)***	(0.022)***	(0.021)**	(0.018)***	(0.026)***
4 Years	0.156	0.157	0.207	0.090	0.121	0.291
	(0.028)***	(0.025)***	(0.030)***	(0.028)**	(0.025)***	(0.036)***
5 Years	0.199	0.200	0.265	0.114	0.154	0.376
	(0.037)***	(0.033)***	(0.040)***	(0.036)**	(0.032)***	(0.049)***
Observations	15,175	33,693	8,744	15,175	33,693	8,744

Note: Standard Errors in Parenthesis;
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

The Less than High School group includes workers with less than 12 years of reported education in RAIS and in the UI database. Similarly, we consider those who report 12 years of completed education as High School Graduates and include workers with more than 12 years of education in the "More than High School Group". The heterogeneous returns are estimated separately for each group, following the control function approach described in Section 3.

Table A1: First Stage Estimation in Brazil and Chile

	Brazil			Chile		
	(1) Experience	(2) Participation	(3) Tenure	(4) Experience	(5) Participation	(6) Tenure
Age	0.674***	-0.542***	0.379***	0.028	0.165	0.250***
	(0.052)	(0.160)	(0.020)	(0.063)	(0.212)	(0.027)
c_{it}	0.009	-0.503***	-0.013	-0.028	-0.031	-0.005
	(0.013)	(0.031)	(0.011)	(0.022)	(0.071)	(0.020)
Age=22	-0.021	-0.189	-0.121	0.104	0.612	-0.241
	(0.165)	(0.433)	(0.132)	(0.226)	(0.741)	(0.199)
Age=23	-0.210	-0.398	-0.273*	0.230	0.209	-0.378*
	(0.181)	(0.496)	(0.120)	(0.220)	(0.727)	(0.172)
Age=24	-0.551**	0.617	-0.585***	0.617**	0.561	-0.444**
	(0.204)	(0.594)	(0.113)	(0.238)	(0.794)	(0.153)
Age=25	-0.564*	0.624	-0.831***	0.830**	0.508	-0.508***
	(0.241)	(0.721)	(0.111)	(0.270)	(0.909)	(0.138)
Age=26	-0.207	1.716*	-1.242***	0.933**	-0.302	-0.571***
	(0.290)	(0.873)	(0.115)	(0.314)	(1.060)	(0.129)
Age=27	-0.043	3.068**	-1.768***	1.582***	0.408	-0.633***
	(0.343)	(1.038)	(0.125)	(0.363)	(1.228)	(0.125)
Age=28	2.022***	0.933	-0.741***	1.741***	0.598	-0.719***
	(0.381)	(1.171)	(0.133)	(0.418)	(1.415)	(0.129)
Age=29				1.610***	-0.396	-0.429**
				(0.488)	(1.652)	(0.142)
Age=22 $\times c_{it}$	-0.020	0.086*	-0.005	0.010	0.011	0.001
	(0.015)	(0.036)	(0.013)	(0.023)	(0.074)	(0.021)
Age=23 $\times c_{it}$	-0.021	0.162***	-0.007	0.013	0.037	0.000
	(0.015)	(0.036)	(0.012)	(0.022)	(0.072)	(0.020)
Age=24 $\times c_{it}$	-0.018	0.167***	0.000	0.010	0.013	-0.003
	(0.014)	(0.035)	(0.012)	(0.022)	(0.072)	(0.020)
Age=25 $\times c_{it}$	-0.025	0.246***	0.000	0.016	0.007	-0.005
	(0.014)	(0.035)	(0.012)	(0.022)	(0.072)	(0.020)
Age=26 $\times c_{it}$	-0.061***	0.204***	0.016	0.033	0.017	-0.003
	(0.015)	(0.036)	(0.012)	(0.022)	(0.072)	(0.020)
Age=27 $\times c_{it}$	-0.081***	0.192***	0.038**	0.021	-0.022	-0.005
	(0.015)	(0.039)	(0.012)	(0.022)	(0.072)	(0.020)
Age=28 $\times c_{it}$	-0.224***	0.401***	-0.033**	0.038	-0.024	-0.005
	(0.014)	(0.036)	(0.011)	(0.022)	(0.073)	(0.020)
Age=29 $\times c_{it}$	-0.183***	0.405***	-0.047***	0.052*	0.017	-0.012
	(0.023)	(0.068)	(0.012)	(0.023)	(0.076)	(0.020)
Age=30 $\times c_{it}$				0.077*	-0.083	-0.021
				(0.035)	(0.116)	(0.021)
HS Graduate	0.044	-0.168	-0.070***	0.024	0.303***	0.038**
	(0.034)	(0.094)	(0.021)	(0.018)	(0.062)	(0.013)
More than HS	0.109	-0.188	-0.106**	-0.251***	-0.623***	-0.068**
	(0.056)	(0.152)	(0.034)	(0.031)	(0.105)	(0.021)
Male	0.046*	0.375***	-0.038***	0.186***	-0.028	-0.094***
	(0.018)	(0.053)	(0.011)	(0.012)	(0.039)	(0.008)
P-Value	0.000	0.000	0.000	0.000	0.000	0.000
Observations	27975	31932	57615	43355	44411	79175
R-Squared	0.329	0.045	0.166	0.312	0.011	0.154

Note: Standard Errors in Parenthesis; *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

In Table A1, we present the coefficients associated with the excluded age and potential experience instruments, along with their interaction terms for the first stage estimations for experience, participation and tenure in both Brazil and Chile. We also include education dummies (workers with less than high school are the excluded category) and a dummy for gender.

Table A2: Coefficients of Residual Terms in Experience and Tenure Estimates

	Brazil		Chile	
	(1)	(2)	(3)	(4)
VARIABLES	Experience	Tenure	Experience	Tenure
Res. Experience	0.002	-0.012	-0.115***	0.021
	(0.026)	(0.016)	(0.029)	(0.018)
Res. Participation	0.007	-0.003	0.030***	0.001
	(0.005)	(0.004)	(0.005)	(0.004)
Res. Exp.× c_{it}	-0.003**	0.001	0.003**	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)
Res. Exp.× Exp.	0.011	-0.006	0.025***	0.001
	(0.008)	(0.004)	(0.008)	(0.005)
Res. Part.× c_{it}	0.000	0.000	0.002***	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Res. Part.× Exp.	0.006***	0.003	-0.003	0.001
	(0.002)	(0.002)	(0.003)	(0.002)
Res. Exp.× Exp.× c_{it}	-0.001	0.000	-0.001**	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Res. Part.× Exp.× c_{it}	-0.000	-0.000**	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Res. Tenure		0.067**		-0.136***
		(0.032)		(0.037)
Res. Ten.× c_{it}		-0.005**		0.008***
		(0.002)		(0.002)
Res. Ten.× Exp.		-0.030***		0.052***
		(0.010)		(0.014)
Res. Exp.× Ten.		-0.019*		-0.036**
		(0.010)		(0.016)
Res. Exp.× Ten.× c_{it}		0.001**		0.003***
		(0.001)		(0.001)
Res. Part.× Ten.		-0.003		-0.001
		(0.003)		(0.003)
Res. Part.× Ten.× c_{it}		0.000		-0.000
		(0.000)		(0.000)
Res. Ten.× Ten.		0.044***		-0.025**
		(0.008)		(0.010)
Res. Ten.× Ten.× c_{it}		-0.003***		0.001
		(0.000)		(0.001)
Res. Ten.× Exp.× c_{it}		0.002**		-0.004***
		(0.001)		(0.001)
Constant	4.939***	-0.014***	5.814***	-0.011***
	(0.017)	(0.004)	(0.022)	(0.003)
Observations	27,975	57,612	43,145	79,175
R-squared	0.231	0.006	0.212	0.005

Note: Standard Errors in Parenthesis; *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

In Table A2, we present the coefficients on the first stage residuals from participation and experience included in the step in which we estimate the returns to experience, as well as the residuals from the first three stage equations in the step in which we estimate the returns to tenure.