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Job Stability in Developing and Developed Countries: Evidence from Colombia and the United States

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Abstract: This paper presents the first systematic comparison of job stability for developing and developed countries, and begins to analyze why long-lasting jobs are less prevalent in developing countries. It compares both cross-section distributions of current job tenure and estimates of job retention probabilities (derived by following synthetic cohorts over time) between Colombia and the United States, and within Colombia before and after a major liberalization of job security legislation. Jobs for males in private sector wage employment are significantly shorter in Colombia than in the United States. The most striking differences in job retention probabilities are for workers in their first year of tenure. Small increases in these low-tenure job retention rates after the liberalization of Colombian job security legislation are consistent with arguments that the original legislation encouraged "rotation" of low-tenure workers, but the effects are quite small, suggesting that differences in legislation do not explain the entire cross-country retention rate difference. Cross-country job length differences diminish, but remain significant, after controlling simply for cross-country differences in the mix of production activities. Thus even when attempting to meet similar production objectives, Colombian employers seem to employ fewer long-term workers. The evidence is consistent with the hypothesis that employers face greater costs of implementing long-term employment contracts or the work organization practices that such contracts facilitate. The paper discusses possible reasons for such cost differences. Regardless of their source, higher costs of long-term employment contracting in developing countries may help explain why those countries exhibit higher rates of self employment and employment in very small establishments, as well as lower levels and growth rates of labor productivity and GDP per capita.

Key Words: Job tenure, job stability, job security legislation, developing county labor markets

JEL Codes: O12, O15, J63

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I. Introduction

Empirical research on urban labor markets in developing countries has generated a wealth of information about wages over the last three decades, but has remained virtually silent about all other features of employment contracts. One of the most important non-wage dimensions of employment contracts is job length or job stability. Not only is job stability of direct interest to workers and governments, as evidenced by the prevalence and political importance of job security legislation. Long-term employment contracts may also play an important role in facilitating modern production and in increasing labor productivity. Theoretical models describe how employers might use long-term employment contracts (involving either efficiency wages or wages that rise with tenure on the job) to increase productivity by facilitating relationship-specific investments in training or screening (Becker, 1962; Hashimoto, 1981; Stiglitz, 1974), or by improving work incentives (Bulow and Summers, 1986; Lazear, 1981).[1] Contracts designed for these purposes are particularly important for facilitating production in large, modern firms, in which employers have arms' length relationships with their employees (Parsons, 1986).

Theoretical links of long-term employment contracting to productivity and to the mode of production suggest that careful study of long-term employment contracting might shed light on the answer to two questions of obvious importance to understanding developing economies and the process of economic development: Why is labor productivity so much lower in developing countries than in developed countries, even after controlling for differences across countries in physical capital per worker and levels of formal education[2]? And, why is the structure of employment so different

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[1] Employers may also use long-term employment contracts to offer workers income insurance in exchange for lower mean wages (Azariadis, 1975).

[2] See the cross-country productivity accounting exercise in Jones and Hall (1996), which accounts for differences in capital per worker and formal schooling. See also case studies, such as those cited in the introduction to Esfahani and Mookherjee, 1995, which compare worker productivity in establishments producing similar goods, using similar capital-labor ratios and machines, and using workers with similar qualifications.

in developing countries, with much larger fractions of the labor force self employed or employed in very small establishments, and much smaller fractions employed in larger, modern establishments of the sort that are the norm in developed countries[3]? Higher costs of long-term employment contracting in developing countries might help answer both questions.

Anecdotal and case study evidence is consistent with the hypothesis of higher long-term employment contracting costs in developing countries. Many employers in developing countries consider high rates of labor turnover to be an important obstacle to the introduction of high performance work practices like total quality management, statistical process control, just-in-time production, multiskilling, and technologies requiring enterprise training more generally (McDermott, 1994; World Bank, 1997). At the same time, case studies document the existence of firms that apparently are successful in using high wages and attractive nonwage benefits and work environments to reduce turnover (Peak, 1993; Humphrey, 1997). More systematic evidence is required, however, before drawing conclusions about the prevalence of long-lasting jobs and the costs of long-term employment contracting.

This paper presents the first systematic evidence on the prevalence of long-lasting jobs for a developing and a developed country, begins to analyze why long-lasting jobs appear to be less prevalent in the developing country, and makes a case for further research on obstacles to long-term employment contracting in developing countries.[4] The study is made possible by the existence of data for Colombia that are somewhat unusual by developing country standards, in that they come from large, high quality household surveys that contain questions on job tenure. They come from

[3] On the prevalence of self-employment, see Schultz (1990). On the size distribution of wage employment, see Little, et al. (1987).

[4] The paper's focus is on urban, nonagricultural labor markets. Some evidence from small samples and agricultural censuses, relating to the use of long-term or "bonded" labor, shed light on employment contract length in rural areas of developing countries (see Mukherjee and Ray, 1995, for citations), but virtually nothing is known about job length in urban areas. Though urban contracting issues have been studied the least, they are of greatest interest as developing countries look to the future of employment and productivity growth.

surveys similar in many respects to Current Population Survey Supplements containing job tenure questions, allowing careful comparisons to the United States at various levels of disaggregation.

Though no one pair of developing and developed countries is typical of all others, the Colombia-United States comparison is a reasonable starting point for the analysis. As is true of developing and developed countries more generally, labor productivity and the share of the labor force in large, wage-employing establishments are indeed lower in Colombia.[5] Furthermore, finding that jobs in private sector wage employment are shorter in urban Colombia than in the United States is especially interesting, because it is likely that differences would be even greater for other developing-developed country pairs. Though jobs are often long in the United States, they are less stable than in several European countries and Japan (OECD, 1993), and the history of greater macroeconomic stability in Colombia than in the rest of Latin America might also suggest that (all else equal) Colombia would exhibit high job stability relative to its neighbors.[6]

The rest of the paper is organized as follows. Part II introduces the data and basic descriptive statistics, indicating that typical current job tenures in private sector wage employment are much shorter in Colombia than in the United States. As discussed in Part III, further analysis is required before concluding that jobs are shorter in Colombia. Current job tenures (i.e. tenures to date on jobs in progress) may appear shorter in Colombia even if jobs have the same probabilities of lasting various lengths of time in both countries, if employment growth rates are higher in Colombia (and for other reasons). Part III takes two approaches to inferring whether jobs are indeed shorter in Colombia.

[5] Jones and Hall (1996) put Colombia's output per worker at about 1/4 that of the United States. A more dated but more detailed source (Morawetz, 1981) reports, for example, that in clothing factories with comparable technologies, the number of classic jeans produced per machine operator in an 8-hour day was 44 in the U.S., 28-30 in East Asia, and 19-24 in Colombia. Lieberman and Hanna (1992, p.122) report that a particular productivity-enhancing innovation, the use of Japanese labor management practices, has been slow to diffuse in Latin America more generally. Evidence on differences in the structure of employment is presented below.

[6] Of course, all else is not equal, because the details of job security legislation differ across countries in Latin America. As will be discussed, the data suggest that these differences might play a smaller role in explaining job tenure differences across countries than one might have expected.

One approach involves conditioning the current job tenure comparisons so as to reduce the importance of differential employment growth rates, and the other involves estimation of job retention probabilities within synthetic cohorts followed across pairs of repeated cross sections. Both approaches indicate that jobs are indeed shorter in Colombia. Retention rate calculations indicate further that the cross-country differences in current job tenures arise primarily because the retention rates for workers in their first year on the job are much lower in Colombia. That is, there appears to be much more churning of very low-tenure workers in Colombia.

Parts IV and V raise two potential explanations for the cross country differences in job length. Part IV exploits the availability of data for Colombia from a 12 year period, in the middle of which Colombia reformed job security legislation, to assess whether the faster rotation of low tenure workers there might be the effect of counterproductive job security legislation. In Colombia in the 1980s, as in much of Latin America, job security legislation may have created an incentive to fire workers at low tenures, by imposing dismissal costs that become relevant only after a brief trial period and that rise with the worker's tenure on the job. Though such an effect is suggested by the slight increase in job retention rates for low tenure workers following a relaxation of the regulations, the effect does not appear large enough to explain the cross-country differences. Thus it appears that the cross-country differences in job tenure reflect more intrinsic labor market differences.

Part V raises a potential explanation for cross-country differences in job tenures that must be ruled out before concluding that the costs of long-term contracting are higher in Colombia. Long-lasting jobs may be less prevalent in Colombia merely because (for reasons unrelated to employment contracting concerns) Colombia is specialized in production activities in which long-lasting employment relationships are less valuable. Disaggregated comparisons of current job tenure distributions and of job retention probabilities indicate that Colombia is indeed specialized to a greater extent in production activities in which long-lasting jobs are less important in both countries, and one cannot rule out the possibility that further disaggregation would cause cross-country differences in job length to disappear. However, both because substantial cross-country job length differences remain even after controlling for the most obvious differences in production structure, and because the differences in production structure might be the endogenous outcome of higher long-term contracting costs, one cannot reject the hypothesis that long-term employment contracting costs are higher in developing countries. Thus further research on employment contracting problems appears warranted.

Part VI sets the stage for future research, by setting out reasons why the cost of long-term employment contracting may be higher in developing countries relative to developed countries, and Part VII concludes.

II. Data and Descriptive Statistics

The Colombian data derive from the Encuesta Nacional de Hogares (ENH) waves of June 1984, 1986, 1988, 1992, 1994 and 1996. The ENH is a survey of approximately 20,000 urban households, which has been conducted quarterly by Colombia's Departamento Administrativo Nacional de Estadistica (DANE) since the mid-1970s. The selected waves were administered in 10 metropolitan areas,[7] and contained supplements including the question "How long have you worked in this establishment or enterprise?", the answer to which is recorded in years. The answers to such a question allow one to describe the distribution of current job tenure, or tenure to date on jobs currently in progress.

The data for the United States derive from Current Population Survey (CPS) Supplements of January 1987, January 1991 and May 1988, which are similar to the Colombian data in both

[7] These include 7 principal cities (Bogota, Cali, Medellin, Barranquilla, Bucaramanga, Manizales and Pasto) and 3 secondary cities from diverse regions (Cucuta, Pereira and Villavicencio).

sample design and phrasing of the job tenure question.[8] The CPS is administered to a nationally representative sample of approximately 50,000 households, but, for comparability with the Colombian sample, I restrict attention to individuals in Metropolitan Statistical Areas with populations of at least 100,000.[9]

For both countries I limit the sample to males aged 15 to 59 in private sector, nonagricultural wage employment. I excluded individuals under 15 because that is the lowest age at which tenure data are available in all datasets. Individuals are excluded if over age 59, because 60 was the legal retirement age (for males) in Colombia during the period from which the data derive. I focus on males because changes over time in labor force participation rates, which are greater for women than for men, create difficulties for job length comparisons, as discussed in Part III. I eliminate the few urban workers reporting agricultural work, because the determinants of job length for such individuals are likely to be quite different from those for the majority of metropolitan dwellers. I focus on wage employment, both because the meaning of a job tenure variable is unclear for the self-employed and because the potential employment contracting problems motivating the analysis arise only when the worker and employer are distinct. I focus on the private sector, for which profit-maximizing models of employment contract design are most relevant.

Whether using the ENH or the CPS, constructing descriptive statistics that reflect the distribution of tenure in the relevant population requires the application of weights, which reflect the number of individuals in the population represented by each observation in the sample. I adjust the

[8] In January 1987 and 1991 the question was "How long has ... been working continuously for his/her present employer (or as self-employed)?" and the response was recorded in either full years or months (if less than one year). In May 1988 the question was "How many years have you been working for your present employer?" and the answer was recorded in years, with a special code for less than one year. The tenure question was administered only to wage employees in half the sample in May 1988.

[9] The smallest metropolitan area represented in the Colombian sample has a population of approximately 228,000 in 1992. The smallest Metropolitan Statistical Area size code included in the CPS sample pertains to areas with populations of 100,000 to 249,000.

weights provided by DANE and the CPS in two ways. First, I adjust for nonresponse on the tenure question. Second, I average weights across primary sampling units (PSUs) within cells defined by age and education. This embodies the reasonable assumption that tenure distributions do not differ by PSU after conditioning on demographic characteristics, and renders more appealing the use of asymptotic standard errors by increasing sample sizes within the cells across which weights vary. The weights employed, and their implications for the construction of test statistics and standard errors, are discussed in more detail in an appendix available from the author.

For describing the current job tenure distributions, I report 10th, 25th, 50th, 75th, and 90th interpolated quantiles. I use interpolated rather than simple quantiles, because tenure is reported in integer values of years, there are thus many observations with exactly the same tenure value, and simple quantiles are insensitive to differences in such distributions. The interpolated qth quantile, Q_q, is related to the simple quantile (i.e. the integer values of tenure such that a population-weighted fraction q of the observations have values less than or equal to that value), T_q, in the following way (for $T_q>0$):

$$Q_q = T_q - .5 + \frac{q - \sum_{i=1}^{N} w_i I(t_i < T_q)}{\sum_{i=1}^{N} w_i I(t_i = T_q)},$$

where q is a fraction defining the quantile of interest[10], w_i is the normalized weight associated with individual i, normalized so that the weights sum across all observations to one, N is sample size, t_{ij} is the integer tenure value for the ij-th observation, and I( ) is the indicator function taking the value 1 if the expression in parentheses is true and zero otherwise. (When T_q equals zero, the formula is adapted to cause the quantiles to fall in the range of zero to .5.) Such statistics summarize succinctly information not only on the simple quantiles (which can be obtained by rounding the interpolated

[10] In the tables I present 10th, 25th, 50th, 75th and 90th quantiles, which correspond to q's of .10, .25, .50, .75 and .90.

quantile to the nearest integer), but also on how "close" the quantiles are to taking higher or lower values. The interpolated quantile statistics, and the derivation of standard errors for them, is discussed in appendices available from the author.

Table 1 puts the job tenure comparisons into context by providing a description of the basic labor market features experienced by males ages 15-59 in Colombia during June 1988 and in the U.S. during January 1987, while Table 2 provides descriptive statistics on job tenure for males in private sector wage employment in those years. Because tenure distributions can be sensitive to business cycle fluctuations, shifting to the right in recessions as employers fire low-tenure workers, it is useful to pick years in which the two countries were in comparable phases of macroeconomic fluctuations. In the chosen years both economies had been growing for several years, the rate somewhat higher in Colombia, as had been true on average over the last several decades. The year 1988 also precedes major changes in job security legislation in Colombia (discussed further below), and can be thought of as reflecting any long run (steady state) consequences of the old job security regime for the current tenure distribution.

According to Table 1, the Colombian labor force is younger, less educated, more likely to be in the labor force and unemployed, less likely to be in white collar occupations, and much more likely to be self employed. Though the overall shares of male adult workers in manufacturing and services are very similar across countries, the composition of those sectors differs, with Colombian manufacturing workers much more likely to be in textiles manufacturing and food processing, and Colombian service sector workers more likely to be in personal services. Using data from the May 1988 CPS and the June 1988 ENH, one also discovers that the share of workers in establishments with over 10 workers is 80 percent in the United States and 53 percent in Colombia.

Table 2 presents statistics describing the distributions of current job tenure reports among private sector wage employees. Chi-squared tests (discussed in an unpublished appendix) reject the hypothesis of identical distributions (Chi-squared test statistics with five degrees of freedom=1215.9), tenures being shorter in Colombia than in the United States. It is useful to note that the cross-country differences in median tenures for males are larger than the male-female differences within either country (statistics for female workers not shown). The simple descriptive statistics suggest large enough differences in private sector wage employment job tenures to merit further investigation.

III. Cross Country Differences in Completed Job Tenure Distributions

The current job tenure distributions compared above are only imperfect reflections of cross-country differences in "job length," "job stability," or the "importance of long-term jobs." This section defines completed job tenure distributions, the measurement of which would allow us better to assess cross country differences in job length, and develops two approaches to drawing inferences about them.

III.A. Completed Job Tenure Distributions and the Importance of Long-Term Jobs

Rather than agreeing on exact job lengths, most workers and employers have implicit understandings about the conditions under which jobs will come to an end (with either a quit or a dismissal), which we can think of as placing probabilities on the jobs lasting various lengths of time. The probabilities of a particular new job lasting various lengths of time describe what we may call the *distribution of completed job tenure* (i.e. tenure when the job ultimately comes to an end). It is characteristics of these distributions that we would like to compare across countries when comparing job stability or the importance of long-term jobs.

If we could draw a large sample of individuals starting jobs at a given time and could observe how long each job ultimately lasted, and if the distribution of completed job tenures (conditional on observed characteristics) were common to everyone represented in the sample, then the distribution of tenures across individuals in the sample would consistently estimate the distribution of completed job tenure faced by any individual. Unfortunately, the job tenure distributions described in the previous section are *distributions of current or incomplete tenure* (i.e. elapsed tenure on a job in progress on the interview date) rather than completed tenure. More important for the present purposes, they are representative of the population of jobs in progress on the survey date (jobs which started at various times in the past) rather than from a population of jobs starting at one time. This creates two problems for drawing inferences about cross-country differences in job length.

First, current job tenure distributions may differ across countries, even when the distributions of completed job tenure are identical, if employment growth rates differ. If employment is growing more rapidly in one country, workers will tend to be found at lower tenures, and jobs will tend to look shorter there, even if all jobs have the same probability of lasting a long time in both places. Second, median current job tenure may differ across countries, even when the medians of completed job tenure distributions and the employment growth rates are identical, if differences in the variance of completed job tenure cause the biases associated with "length-biased sampling" to be more severe in one country than the other (Salant, 1977). Because long-lasting jobs are more likely than shorter jobs to be sampled at any point in time, samples of current job tenures are biased toward longer tenures,[11] and this bias is more severe the higher is the variance of the underlying distribution of completed job tenures.

The rest of this section makes two attempts to draw inferences about cross country differences in (some features of) completed job tenure distributions. The first attempts to mitigate the effect of differential employment growth rates by conditioning current job tenure comparisons in ways that should render the relevant growth rates more similar, but does not compensate for the effects of length-biased sampling. The second, which involves estimation of job retention rates

[11] While length biased sampling tends to render the mean of current job tenure higher than the mean of completed job tenures, the measurement of incomplete rather than completed tenures tends to render the mean of incompleted job tenures lower. As described in Salant (1977), the effect of length-biased sampling outweighs the effect of truncation when the probability of a job ending, conditional on having lasted to tenure t, falls as t rises.

within synthetic cohorts across cross sections, describes some features of the distribution of completed job tenures directly and mitigates both problems.

II.B Cross Section Current Tenure Distributions Revisited[12,13]

At the aggregate level, employment may be growing more rapidly in one country, tending to make current job tenures there look shorter, for several reasons: more rapid population growth, more rapid increases in labor force participation, or more rapid immigration into the large cities represented by the samples. By restricting attention to males and by conditioning on age as in Table 3, I hope to mitigate the effect of all three sources of differential aggregate growth rates. Conditioning on age, by producing tenure statistics within age groups, is a simple (albeit imperfect) way of controlling for higher population growth rates. Focusing on males mitigates the effects of differential labor force participation rates, which are most important for women. Conditioning tenure comparisons on age and focusing on older age groups should also mitigate the effect of differential immigration on job stability comparisons. Rural-urban migrants in developing countries typically move while young.[14] Thus the effect of their entry into the urban labor force on current tenure distributions should be smaller when looking at older age groups than when looking at younger age groups.[15]

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[12] Some authors use cross section current tenure data to estimate directly job retention probabilities. The method, attributed to Hall(1982) and often called "contemporaneous retention rate calculations," essentially accounts for changes in employment associated with population growth, but assumes away changes in employment arising out of changes in labor force participation or immigration. Such changes may be important and may differ across countries, thus the approach in Section III.C to estimating job retention probabilities is superior.

[13] This section maintains the assumption that completed job tenure distributions have been constant over the relevant past. This assumption, which is made palatable by the observation that the large majority of jobs in both economies began in the previous five years (implying a relatively short "relevant past"), implies that the share of jobs in today's sample that started s years in the past contains information about the probability with which a job starting today will last at least s years.

[14] According to the ENH data, over eighty percent of workers in their 50s had not migrated in the last 10 years.

[15] In another attempt to assess the likely importance of immigration into large cities in Colombian job tenure distributions, I recalculated the Colombian statistics using only observations from "natives," defined as those who have lived in the municipality in which there were interviewed since at least age 14. (Note that if working life for most

Table 3 presents the results of disaggregating the interpolated median tenure comparisons by age group. The cross country differences in private sector wage employment remain striking. By the time private sector wage employees have reached their 50s, they enjoy interpolated median tenure of nearly 14 years in the U.S., but only 8.5 years in Colombia. Chi-squared tests of the hypothesis that tenure distributions are identical across countries reject at the .001 significance level for all age groups but the youngest.

Consideration of possible biases associated with more disaggregated employment growth rates suggests that Table 3 may in fact understate cross-country differences in job tenure. Because Table 3 focuses on private sector wage employment, it may yield misleading cross-country comparisons of job length, even when differences in aggregate employment growth rates have been controlled, because private sector wage employment may be increasing or decreasing more rapidly as a share of total employment in one country than another. In addition, because Table 3 disaggregates by age within private sector wage employment, relevant private sector wage employment growth rates may differ across countries if individuals have a greater tendency to move out of wage employment into self-employment as they age in one country. Figure 1, which plots the percentage of males in private sector wage employment by age group, for six Colombian and two U.S. datasets, indicates that while age specific private sector employment rates are not changing much over time within countries, they do decline more rapidly with age in Colombia than in the United States. The decline is likely to be the result of true life cycle movements out of wage employment into self employment, and not just the result of older cohorts having higher self-employment rates throughout their lives, because self-

people begins after 14 years of age, then higher immigration by individuals who are less than 14 years old does not lead to spuriously lower job tenures among workers.) This is a very narrow definition of nonmigrant, because it excludes some workers who have merely moved within the metropolitan area covered by the sample. Somewhat surprisingly, rather than increasing tenures, this produced shorter tenure statistics. This suggests that the behavior of migrants and nonmigrants differ in some important ways that might merit further research. For the purposes of the current study, however, I chose to produce statistics for the population at large rather than for the apparently less stable nonmigrant sample.

employment rates within age groups do not fall over time.[16] Thus it seems likely that employment growth rates in private sector wage employment among older workers are higher in the U.S. than in Colombia, in which case current job tenure comparisons for that group understate the extent to which jobs are less stable in Colombia.

III.C Synthetic Cohort Job Retention Rates

A method of comparing features of completed job tenure distributions that circumvents problems related to differential employment growth rates and length-biased sampling involves the estimation of job retention probabilities within synthetic cohorts. *X-year job retention probabilities* can be estimated if cross sections for two years that are X years apart are available, and if the application of sample weights allows reasonably precise estimation of the population of workers at each tenure in each year. They are derived simply by dividing the estimated population of individuals with tenure T+X and age A+X in the second year by the estimated population with tenure T and age A in the first year. Derivation of standard errors that account for the weighting required is discussed in an appendix. Differences across countries in job growth rates do not muddy these estimated retention rates, because the probabilities are derived by following sets of jobs over time, rather than by comparing numbers of jobs that started at different times and survived until a specific interview date. The effects of length-biased sampling are also largely eliminated.[17,18]

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[16] Additional evidence that this reflects true life cycle movement may be drawn from the Colombian data, in which many older self-employed workers report having started their current self-employment spell in the last few years, and in which many of the older self employment job starters, many report that their previous job was in wage employment.

[17] They are not eliminated entirely, because rather than observing job retention rates for all jobs starting in a given year, we observe all jobs that started in a given year and lasted at least until the interview date. Thus very short jobs may be underrepresented (to differing degrees in the two countries).

[18] The only potential drawback to using such estimates is that they are more demanding of the population weights provided by DANE and the CPS than are comparisons of single cross-section tenure distributions. In single cross section comparisons it is vital only that the relative weights placed on individual observations be accurate. When calculating retention rates, it is also important that the absolute level of the weights be correct in both years. If, for example, the weights in the second year are too small, leading to the underestimate of the population at any given

It is useful to allow such retention rates to differ by tenure and age in the first year. Retention rates are likely to rise with initial tenure, at least after a fall over the first few months of tenure on the job (Farber, 1994), because higher tenure workers have made larger investments in assets specific to the job, because workers who remain until higher tenures are a select group whose characteristics render them less likely to quit or be fired, or because, as is the case in Colombia, regulations on dismissals cause the cost of dismissal to rise with job tenure. Retention rates may also vary with age, if, for example, workers' abilities to identify jobs that are good matches improves as they age.

Data availability restricts me to comparing 4-year job retention rates. Table 4 presents job retention rates in private sector wage employment over the intervals 1984-88 and 1988-92 in pre-reform Colombia and 1987-91 in the United States. The estimated standard errors rise with initial tenure, but are all quite small. The point estimates for Colombia that inspire the greatest confidence are for the lowest three initial tenure categories, which are quite stable across the two periods for which estimates are possible. Informal calibration of job survival functions based on these job retention probabilities suggests that these differences in job retention rates imply large differences in typical job length, and are consistent with the cross-section current job tenure distributions (details available from the author). The most striking differences in job retention rates are for workers in their first year on the job (that is, with initial tenure equal to zero), who are more than one and a half times as likely to retain their job in the United States as in Colombia. The figures seem to suggest that lower current tenures in Colombia arise primarily as a result of more rapid rotation of low tenure workers. At higher tenures retention rates are similar across countries or even higher in Colombia.

Table 5 disaggregates the retention rate comparison for private sector wage employees by initial age category as well as years of initial tenure. The estimates suggest that the lower job

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tenure level, then all job retention rates will look too small. Fortunately, DANE and the Department of Planning undertook a major revision of the Colombian weights in 1993, with the aim of reconciling them with independent population estimates and making them consistent across years.

retention probabilities at low tenures in Colombia are experienced by workers of all ages, with cross-country differences somewhat higher in higher age groups.

IV. Job Security Legislation Reform and Job Stability in Colombia

A first potential explanation for the cross-country differences in job length pertains to differences in job security legislation. The pattern of much lower job retention rates in Colombia for workers in their first year of tenure on the job, and higher retention rates for high tenure workers, is consistent with the stories often told about the counterproductive effect of job security legislation in Latin America. In Colombia in the 1980s, as in much of Latin America, job security legislation rendered dismissals very costly while also creating incentives to fire workers at low tenures (Colombia, 1988; Lopez Castano, 1994). Dismissals were costly, because (after a two month trial period) individual workers fired without "just cause" were entitled to severance payments, after 10 years on the job individual workers fired without just cause could sue for large indemnities and reinstatement in the old job, and group layoffs were made difficult by requirements for approval by the government. The legislation created some incentive to fire workers early, however, for several reasons. Workers in their first two months of tenure on the job were in a probationary period, during which they could be fired at will. After their workers had acquired at least two months' tenure, employers had to set aside sums each year equal to the additional severance payments to which workers became entitled, and these sums increase with tenure for two reasons. For workers with 1 to 5 years of tenure, the incremental payment liabilities associated with an additional year of tenure were equal to 15 days pay, while for workers with 6 to 10 years tenure they were equal to 30 days pay. On top of this, liabilities associated with any year of tenure on the job were calculated with respect to average pay in the last year on the job; this "retroactivity" provision meant that for workers who received raises, sums equal to the raise times all days pay previously deposited in the accounts had to be added to the accounts. Incentives to fire workers at early tenure were also created by provisions giving workers with 10 years tenure the right to receive a pension at age 60, and workers with 15 years tenure the right to receive a pension at age 50. The net effect of the legislation is widely believed to be the nearly perfect stability of workers with more than ten years tenure, but the customary firing of workers before they reach ten years tenure, and perhaps a tendency to "rotate" very low tenure workers.

Availability of Colombian data from 1984 through 1996 provides an opportunity to assess the likely importance of Colombian job security legislation in explaining the cross-country differences in job retention rates for low tenure workers. The years 1990 and 1991 were watershed years in which the Colombian economy opened and liberalized along many dimensions, including a liberalization of job security legislation. In the rhetoric of the reform, the revision of job security legislation was intended to render the economy more flexible and thus more able to compete in the world economy. At the same time, economists expected it to increase job stability by reducing the incentives to fire workers at early tenures. The key provisions included a widening of the legal definition of "just cause," the imposition of penalties on bureaucrats who did not process requests for mass layoffs quickly, elimination of the retroactivity provision (though this was partially compensated by an increase in the number of days pay to be set aside each year), elimination of the ability of workers with more than ten years tenure to sue for reinstatement, and elimination of the rights to pensions at age 60 and 50 for workers with 10 and 15 years of tenure. Workers earning more than 10 minimum wages were also newly allowed to opt to receive higher wages ("salario integral") in lieu of severance pay and other benefits.

Because the new provisions applied only to jobs starting after the law's implementation and to other workers opting to switch to the new regime, the full effect of the legislative changes is unlikely to be observed in the data available, which extend only until five and a half years after the new law was put into effect. Fortunately, the workers whose retention rates we would most like to observe under both old and new regimes are workers in their first year on the job. The data allow assessment of two- and four-year job retention probabilities for such workers under both old and new regimes. The four-year spans 1984-88 and 1988-92 and the two-year spans 1984-86 and 1986-88 describe job stability before the legislation change, and the spans 1992-1996, 1992-94 and 1994-96 describe stability after. Though at higher tenures even the retention rates for the latter spans may pertain largely to workers still under the old regime, workers in their first year of tenure on the job in 1992 and 1994 should be under the new regime.

Tables 6 and 7 indicate four- and two-year Colombian job retention rates for the pre- and post-reform intervals.[19] As shown in Table 6, four-year job retention probabilities in private sector wage employment make a marginally statistically significant increase from the pre-reform to the post-reform period, but the effect is small compared to the size of cross-country differences in these rates. The two-year retention rates shown in Table 7 suggest that low-tenure job retention rates might be inching up over the years following the reform (perhaps as employers learn new ways to organize labor that take advantage of the new job security regime), but again, the changes are very small.

Given that 1990 and 1991 brought liberalization of policies regarding trade, foreign exchange and foreign investment as well as dismissals, it remains possible that the liberalization of dismissal regulations alone would have increased job retention rates for workers in their first year on the job even more, but that exposure to increased international competition caused dislocation at all initital tenures, preventing a larger increase in those rates. Though this cannot be ruled out, the effects of increased competition seem unlikely to mask very large labor legislation effects. Both the initial slowdown and the subsequent acceleration of growth associated with the liberalization in Colombia were surprisingly mild. Official statistics (available at http://www.dane.gov.co) indicate that while

[19] Interpolated quantiles of the current job tenure distribution show very little change over 6 cross sections. Job retention probabilities for low tenure workers should be more sensitive to changes in the legislation.

the growth rate of real GDP fell from 3.4 percent in 1989 and 4.3 percent in 1990 to 2.0 percent in 1991, it rose to 4.0 in 1992, and 5.4 and 5.8 in 1993 and 1994. Furthermore, there was no surge either in official unemployment or in labor force nonparticipation.

The results of this section suggest that Colombian job security legislation may indeed have been partially counterproductive, contributing to the higher separation rate among low tenure workers in Colombia relative to the United States, but that it at best explains only a small fraction of the total U.S.-Colombia difference in separation rates. Thus while job security legislation has some role in shaping the environment within which workers and employers make choices about employment contracts, quits and dismissals, we must look to other features of the economic and institutional environment, such as those discussed in the next two sections, if we are to understand why jobs are shorter in Colombia than in the United States.

V. Specialization and the Importance of Long-Term Jobs

A second potential explanation for the lower incidence of long-term jobs in Colombia relative to the United States has to do with developing countries' greater specialization in production activities with simpler technologies and smaller scale:[20] long-lasting jobs may be less important in developing countries simply because fewer workers there are involved in production activities in which long-term employment arrangements (and the work organization practices they facilitate) are valuable. If developing countries are specialized in production activities in which jobs are easily learned and in which difficult-to-measure worker qualities have little effect on productivity, then there may be few training and screening costs on which to economize, and little benefit to designing contracts that reduce turnover. Similarly, they may be specialized in production activities in which the small size of the organization renders it cheap to monitor and supervise workers directly. In this

[20] See footnote 3.

case there may be little benefit to designing high-wage contracts that economize on monitoring costs.

Such an explanation would have to be ruled out as the entire explanation for cross-country differences in job length before concluding that the cross-country differences length provide evidence of higher costs of long-term employment contracting in developing countries. With this is mind, this section makes a crude attempt to estimate cross-country differences in job tenure distributions that control for differences across countries in the composition of production activities. Turning first to the current job tenure distribution comparison, I proceed as follows. I use CPS data from May 1988, which includes an indicator of establishment size, and the June 1988 ENH. I divide the samples into cells defined by eight productive sectors (food, beverage and textile manufacturing; other manufacturing; construction; commerce; transportation and communication; finance and other business service; personal services and other services), two broad occupation groups (blue and white collar) and two establishment sizes (over and under 10 employees). The broad occupation control is introduced to deal with the concern that even within industries, developed country establishments specialize more in higher level segments of industrial activities. Controlling for establishment size may improve controls for the sorts of goods and services being produced, as well as controlling crudely for capital intensity and organization size.[21] Given the concerns of Part III, I disaggregate also by age. I calculate interpolated medians within cells, and regress them on cell characteristics. Following Farber (1995), I weight the regressions by cell size. I exclude cells with fewer than 20 observations and I report robust standard errors. I interact age categories with all other cell characteristics, allowing people to "accumulate" tenure as they age at different rates in different countries and in different types of production activities.

The results are shown in Table 8. The coefficient on the interaction between the US indicator

[21] The distinction between establishments with more and fewer than 10 employees is the only size distinction I can make in both Colombian and U.S. datasets. Colombian surveyors distinguish 4 sizes, of which "over 10" is the largest, while CPS surveyors distinguish 6 sizes, of which less than 10 is the smallest.

and a specific age category indicator estimates the U.S.-Colombia difference in interpolated median tenure within that age category. The interpretation of other interaction terms is similar. In order to facilitate comparison with the retention rate regressions that follow (in which it is impossible to control for establishment size), I report results in which I both do (column 1) and do not (column 2) control for establishment size. Many standard errors are large, and tenure differences associated with some cell characteristics do not rise monotonically with age, thus the results do not appear as strong as those of previous sections. They nonetheless suggest the following. Controlling for production activity diminishes cross-country differences in interpolated medians (compared to the figures in Table 3), but does not cause them to disappear. Tenures remain statistically significantly different across countries. With the overall Colombian median tenure for the oldest age brackets at 5 to 8 years, the 1- to 2-year cross-country difference also appears economically significant. Large establishments have much longer tenures than smaller establishments, but adding the size control reduces the cross-country tenure differences only modestly.

Turning next to the job retention probability comparison, I proceed as follows. I divide the samples into cells defined by the eight productive sectors and two broad occupation groups (blue and white collar). I disaggregate also by initial tenure categories, and interact the initial tenure indicators with other cell characteristics, allowing producers in one country or in some production activities to rotate low tenure workers more while retaining high-tenure workers with higher probability. I calculate job retention probabilities within cells, and regress them on cell characteristics, again weighting the regressions by cell size, excluding cells with fewer than 20 observations, and reporting robust standard errors.

The results are shown Table 9. The coefficient on the interaction between the US indicator and a specific initial tenure category indicator estimates the U.S.-Colombia difference in job retention rates within that initial tenure category, where a difference of .10 indicates a 10 percentage point difference. Industry and occupation controls are important in explaining job retention probabilities. But even after controlling for industry and occupation, the 4-year job retention probability for workers in their first year of tenure on the job remains 11 percentage points higher in the U.S. relative to Colombia. This difference is nearly as large as the "raw" difference shown in Table 4. Given the small effect that inclusion of size controls had on cross-country differences in interpolated median tenures, it seems unlikely that this large difference in low-tenure job retention probabilities would disappear if size controls could be added.

The results indicate that Colombia in indeed specialized to a greater extent in activities in which jobs tend to be shorter in both countries, but that even after controlling for these differences in production mix, jobs continue to look shorter in Colombia than in the United States. As in the more aggregated results, the most striking cross-country difference is in job retention probabilities for workers in their first year on the job.

Unfortunately, such results rule neither in nor out the possibility that the lower incidence of long-term jobs in Colombia is entirely the result of exogenously greater specialization in technologies requiring few long-term employment contracts. On the one hand, it remains possible that with more detailed controls for production activities the cross-country tenure differences would disappear. On the other hand, even if the cross-country job length differences had disappeared after controlling for differences in production mix, it remains possible that Colombia's specialization in establishments with smaller scale and simpler technologies is in part the endogeous outcome of higher costs of long-term employment contracting costs. Thus, though explanations for cross-country differences in job length based on exogenous differences in production mix cannot be ruled out, neither can the hypothesis that the costs of long-term employment contracting inhibit development in poor countries.

VI. Possible Sources of Higher Long-Term Employment Contracting Costs

Reflection on theoretical models of long-term employment contracting produces four sets of potential explanations for higher long-term employment contracting costs. The first set applies to long-term employment contracts across the board: the effective rate at which workers discount future wages on the current job might be higher in developing countries, both because the rate at which jobs end for reasons exogenous to contracting considerations might be higher, and because credit constraints on workers might be more severe at lower income levels and in more fragmented credit markets. In both market-clearing and efficiency wage versions of models in which promises of high future wages are used to reduce turnover or shirking, an increase in this rate increases the relative cost of hiring workers under effective long-term contracts.[22] Higher exogenous separation rates might result from less stable macroeconomic policies, greater sensitivity of economies to fluctuations in agricultural productivity or world commodity markets, poor infrastructure and communication systems that make it difficult to maintain relationships with suppliers and customers, all of which might produce greater "churning" of firms. Indeed, Roberts and Tybout (1996) report greater volatility in employment associated with expansion, contraction, births and deaths of firms in several developing countries relative to developed countries.[23] Poorer public health and health care delivery systems may also exogenously increase separation rates by increasing the probability that workers must quit to care for their own health or that of the their relatives.

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[22] See Esfahani and Mookherjee (1995) for a specific model in which increases in the effective discount rate lead to reductions in the share of workers employed under contracts paying premium wages designed to reduce shirking.

[23] In particular, using firm-level panel data, Roberts and Tybout (1996) report rates of job creation (the number of jobs "gained" at firms that came into existence or increased the total number of employees between years of the panel, as a percentage of the total number of jobs in the initial year) and job destruction (the number of jobs "lost" at firms that went out of business or reduced the total number of employees between years of the panel, as a percentage of the total number of jobs in the initial year) that are higher in Colombia (in the mid-1970s through the mid-1980s) than in the U.S. and Canada (Figure 1.1, p.4). Whether the higher probabilities of firm failures or contractions plays an important role in inhibiting the development of productivity-enhancing long-term employment contracts is an open question.

A second set of reasons for higher costs of long-term employment contracting are specific to long-term employment contracts used to enhance productivity in specific ways. For example, if workers with lower quality and quantity of formal schooling are more costly to train, then long-term contracts used to facilitate investments in on-the-job training will be less attractive where formal education outcomes are inferior.[24] Esfahani and Mookherjee (1995) provide another possible example: under their assumptions regarding preferences toward hard work, the premiums (relative to the cost of spot market labor) required to induce workers not to shirk are higher where income levels are lower, thus long-term contracts used to provide good work incentives should be at a particular disadvantage in developing countries.

A third set of reasons for higher costs of long-term employment contracting costs in developing countries relates to more severe financial constraints, which would inhibit long-term employment contracts used to facilitate any relationship-specific investments with up-front costs paid in part or in full by the employer, such as investments in screening, specific training, and even general training in the presence of information problems (Acemoglu and Pischke, 1996). Finally, differences in the use of long-term contracting may represent multiple equilibria in the cost and incidence of long-term contracting. For example, Chang and Wang (1995) point to the possibility of self-fulfilling expectations about the average quality of workers who leave jobs to look for new ones.

Identifying which, if any, of these explanations play important roles in explaining why job tenures are shorter in Colombia than in the United States, and perhaps in developing relative to developed countries more generally, would enhance our understanding of obstacles to economic development.

[24] U.S.-Colombia differences in job tenure distributions decrease further when, in exercises like those of the previous section, controls are added for differences in shares of the population having completed secondary education. Unfortunately, at this level of disaggregation, statistical results appear less reliable, thus they are not presented in the tables.

VII. Conclusion

This paper shows that the typical worker in Colombia has been on the job for a substantially shorter period of time than his counterpart in the United States. Such cross section differences in job tenure are not a mere artefact of higher employment growth rates in Colombia; they indeed reflect that significantly smaller fractions of jobs are long-lasting there. The lower prevalence of long-lasting jobs in Colombia appears not to be the result of counterproductive job security legislation, nor does it seem likely to be the result of exogenously greater specialization in production activities in which there is less need for long-lasting employment relationships. Thus the evidence in the paper is consistent with the hypothesis that the cost of long-term employment contracting is higher in Colombia relative to the United States, and perhaps in developing relative to developed countries more generally. It provides a motivation for future research regarding the reasons why such costs might be higher in developing countries. Because long-term employment contracts might be important for facilitating a variety of productivity-enhancing labor practices, and because they may be especially important for doing this in larger establishments in which employers have arms' length relationships with their workers, the evidence also suggests the importance of future research on the extent to which higher costs of long-term employment contracting help explain why labor productivity is lower, and why the share of workers either self employed or employed in very small establishments is higher, in developing countries.

The results also call for a rethinking of the recent literature on cross-country differences in growth rates of labor productivity and GNP per capita. At the very least, the descriptive statistics of Part II indicate that standard measures of the stock of formal schooling could be highly inadequate measures of the "human capital" thought so important to the growth story. Job-specific experience may vary greatly across countries, in a way that has so far been ignored in the empirical cross-country growth literature (following Barro, 1991).

More important, a rethinking of theoretical models of cross-country growth differences (such as those following Lucas, 1988) may be required. Theoretical models of cross-country differences in productivity and growth have tended to abstract from employment contracting difficulties, assuming that labor with a given level of formal schooling is a homogeneous commodity that may be bought and sold in well-functioning spot markets free of information problems. They direct policymakers to be concerned about formal schooling levels, incentives toward R&D and technology transfer, and to the macro level institutions defining the scope for diversion by producers or government officials. The evidence presented above motivates interest in a new generation of models focussing on the micro level institutions through which workers are brought into the production process. Obstacles to the development of long-term employment contracting might inhibit growth, both because long-term contracts may provide workers with incentives to share information that leads to productivity advance (as in Levine and Parkin, 1993, p.253-3), and because productivity may tend to advance more rapidly in the larger, modern establishments that make intensive use of long-term employment contracts. Thus, where the macroeconomic environment is less stable, where the cost of providing workers with a given set of skills through on-the-job training is higher, or where the costs of long-term employment contracting are higher for other reasons, employers might rationally make fewer investments in the development of productivity-enhancing long-term employment relationships, with potentially important ramifications for both the level and growth rate of labor productivity.

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Appendix 1
Interpolated Quantiles

When describing current job tenure distributions, I use interpolated (Farber, 1995) rather than simple quantiles, because tenure in years is reported in integer values, and thus contains many observations with exactly the same reported value. Simple quantiles can be quite insensitive to shifts in such distributions. For example, in each of the following two samples the sample median is 2, though there is a sense in which the sample median is "just 2" in the first and "almost 3" in the second.

Tenure observations in Sample A: 1 1 1 1 2 2 2 2 3 3

Tenure observations in Sample B: 1 1 2 2 2 2 3 3 3 3

If true job tenures are continuous but are rounded to integer reports in years, then one might guess that the true mean is closer to 1.5 in the population corresponding to Sample A and closer to 2.5 in the case of Sample B. The median appears closer to 3 in sample B because a larger fraction of the observations with the value 2 (the integer quantile) must be included with the observations with value 1 (all values lower than the integer quantile) to achieve 50 percent of the observations. That is, the 50 percent line is farther to the right in the list of observations with the value 2.

The interpolated quantile Q_q is related to the simple quantile T_q (i.e. the integer value of tenure such that a population-weighted fraction q of the observations have values less than or equal to that value), in the following way (for $T_q>0$):

$$Q_q = T_q - .5 + \frac{q - \sum_{i=1}^{N} w_i I(t_i < T_q)}{\sum_{i=1}^{N} w_i I(t_i = T_q)},$$

where q is a fraction defining the quantile of interest[1], w_i is the normalized weight associated with individual i, normalized so that the weights sum across all observations to one, N is sample size, t_{ij} is the integer tenure value for the ij-th observation, and I( ) is the indicator function taking the value

[1] In the tables I present 10th, 25th, 50th, 75th and 90th quantiles, which correspond to q's of .10, .25, .50, .75 and .90.

1 if the expression in parentheses is true and zero otherwise. (When T_q equals zero, the formula is adapted to cause the quantiles to fall in the range of zero to .5.) If one conditions on T_q (as suggested by Farber, 1995), one may derive standard errors by applying the delta method, as discussed in Appendix 2.

If the sample were self-weighting, so that w_i=1 for all i, the formula for the interpolated quantile could be reexpressed in a somewhat more intuitive way, as

$$Q_q = T_q - .5 + \frac{q - Fr(T<T_q)}{Fr(T=T_q)},$$

where Fr( ) indicates the fraction of the sample satisfying the inequality or equality in parentheses. According to this formula, if the integer quantile is 2, one calculates the interpolated quantile by subtracting .5 from 2 (so that the interpolated quantile lies between 1.5 and 2.5), and then adding the fraction of the observations with value 2 that must be included with observations of lower value to achieve a fraction q of the entire sample.

Interpolated quantile reports allow one to recover the integer quantiles (because, for example, interpolated quantiles in the range [1.5,2.5) must correspond to an integer quantile of 2), but also summarize additional information on how “close” the quantile comes to taking a higher or lower integer value. The interpolated quantiles would provide consistent estimates of the true, continuous quantiles if the true tenures were uniformly distributed in the range [T-.5,T+.5). But because this assumption is unlikely to be a good approximation to reality, it seems preferable to think of the interpolated quantiles merely as convenient ways to summarize information about the discrete distribution.

Appendix 2

Weighting, Standard Errors and Testing

Development of standard errors and test statistics is complicated somewhat by the need to employ weights for describing population distributions. Weights are employed because tenure distributions are suspected to differ systematically across groups whose proportions in the sample differ from their proportions in the population. In deriving test statistics and standard errors, I must acknowledge explicitly these potential differences in group-specific distributions.

Both the ENH and the CPS employ stratified random samples of Primary Sampling Units (PSUs), which are small neighborhoods containing on the order of 10 to 20 households, all of which are typically interviewed. In principle the survey is self-weighting, in the sense that each stratum is sampled in proportion to its share in the population. In practice, however, both DANE and the CPS provide weights that adjust for three potential departures from a self-weighting sample: (1) When interviewers discovered more than 20 households in a PSU, they followed rules for selecting a random subsample of the PSU's households. The weights for those households are inflated accordingly. (2) Some households refused to respond. (3) Errors in either the sampling frame or the sampling process led to divergences between sample-based population estimates and independent population estimates within demographic groups. The weights that adjust for all three potential problems vary across demographic groups (defined by gender, age, and, in the case of the U.S., race) and PSUs.

I made two adjustments to these weights. First, because I am interested in comparing population tenure distributions across countries and because I suspect that tenure distributions differ significantly by age, education and job type (private sector wage employment, public sector wage employment, self employment), I adjusted the weights for nonresponse to the job tenure question, multiplying them by the inverse of the response rate among workers within age-education-job type categories. Tenure nonresponse rates were quite low (on the order of five percent or less) for the Colombian datasets and for the U.S. 1987 and 1991 datasets, but were much higher (on the order of 20 percent) for the U.S. 1988 dataset. After I make this adjustment, the weights vary by education level and job category, as well as demographic groups and PSUs.

Second, because I am willing to assume that tenure distributions are the same across PSUs (after conditioning on demographic characteristics, education level and job category), I average the weights across PSUs within age-education-job category groups.[2] This assumption makes the asymptotic standard errors derived below more appealing as an approximation. As will be seen, once I account explicitly for the role weights play in the analysis, I must employ asymptotics in which sample size gets large within each group for which weights differ. When weights are allowed to differ across groups defined by PSUs as well as by age, education and job category, the groups are very small, making it unlikely that the asymptotics provide a good approximation. Once weights are averaged as suggested here, the group-specific sample sizes are large enough to make the asymptotic results more appealing.

The weights provided by the survey organization indicate the number of individuals in the population represented by the individuals in the sample. After making the two adjustments just described, I divide them by total population size, so that they sum across all observations in the sample to 1. The normalized weights represent the ratio of the group's share in the population (π_j) to the group's sample size (n_j), where $j=1,\ldots,J$ indexes the groups across which weights and tenure distributions are allowed to differ. When considering the asymptotic properties of the statistics described below, I assume that group definitions and the percentage distribution of sample observations across groups remain fixed as sample size (N) becomes large. This implies that the J group-specific sample sizes, $n_1, \ldots, n_J$ all become large as N becomes large.

Chi-Squared Test of Equality of Current Tenure Distributions. The aggregate distribution we attempt to estimate when producing descriptive statistics on current job tenure distributions is a weighted average of group-specific distributions, with weights given by $\pi_1,\ldots,\pi_J$. I characterize this distribution by the probabilities it attaches to a randomly chosen tenure (τ) falling in six of the seven "bins" indicated in Table 2, when the tenure is drawn from a mixture of group-specific distributions in which the distributions enter with weights $\pi_1,\ldots,\pi_J$. (Because probabilities of mutually exclusive and exhaustive events must sum to one, six probabilities are sufficient for describing the distribution across seven bins.) This characterization of the distribution may be written as the vector

[2] Note that it is unnecessary to use weights for describing population distributions, even when sampling ratios vary across segments of the population, as long as the distribution is identical across segments.

$$F = \begin{bmatrix} Pr(\tau \in bin_1) \\ \dots \\ Pr(\tau \in bin_6) \end{bmatrix} = \sum_{j=1}^{J} \pi_j F_j = \sum_{j=1}^{J} \pi_j \begin{bmatrix} Pr_j(\tau \in bin_1) \\ \dots \\ Pr_j(\tau \in bin_6) \end{bmatrix}.$$

F may be estimated in an unbiased fashion by

$$\hat{F} = \sum_{j=1}^{J} \pi_j \hat{F}_j = \sum_{j=1}^{J} \pi_j \frac{1}{n_j} \sum_{i=1}^{n_j} \begin{bmatrix} I(\tau_{ij} \in bin_1) \\ \dots \\ I(\tau_{ij} \in bin_g) \end{bmatrix} = \sum_{j=1}^{J} \sum_{i=1}^{n_j} w_j \begin{bmatrix} I(\tau_{ij} \in bin_1) \\ \dots \\ I(\tau_{ij} \in bin_6) \end{bmatrix},$$

where I(.) is an indicator function taking the value 1 if the expression in parentheses is true and zero otherwise, and τ_{ij} is the value of tenure for the i,jth observation. Letting $F_j(t)$ denote the t-th element of the vector F_j, the random vector

$$x_{ij} = \begin{bmatrix} I(\tau_{ij} \in bin_1) \\ \dots \\ I(\tau_{ij} \in bin_6) \end{bmatrix}$$

has mean F_j and covariance matrix Q_j, which has diagonal elements $F_j(t)[1-F_j(t)]$, t=1,...,6, and off-diagonal elements $-F_j(t)F_j(t^*)$, $t \neq t^*$. $\hat{F}_j$ is the sample mean of x_{ij} within group j. By the Central Limit Theorem, as $n_j \to \infty$,

$$\sqrt{n_j}(\hat{F}_j - F_j) \xrightarrow{d} N(0, Q_j).$$

Applying the rules of convergence in distribution, $\hat{F}$, too, is asymptotically normally distributed, and has mean equal to F. Because group-specific statistics are independent, $\hat{F}$ has asymptotic covariance matrix given by

$$Q = \sum_{j=1}^{J} \pi_j^2 \frac{1}{n_j} Q_j .$$

We are interested in comparing the distribution of F to an independent distribution G from another country, which is estimated in similar fashion, and has comparable asymptotic covariance matrix Q'. Again invoking rules of convergence in distribution, and under the null hypothesis that F=G, we have the result that the statistic

$$[\hat{F} - \hat{G}]^T [Q + Q']^{-1} [\hat{F} - \hat{G}]$$

is asymptotically Chi-squared with 6 degrees of freedom.

The asymptotic covariance matrix Q may be estimated in the following way. Within each group, relative frequencies of observations in each bin can be used to estimate $F_j(1),\ldots,F_j(6)$. With these probability estimates, one can construct estimates of Q_j, which can be weighted by $\pi_j w_j$ and summed across groups to produce an estimate of Q. Similar derivations produce estimates of Q'.

<u>Standard Errors for Interpolated Quantiles.</u> The qth interpolated quantile Q_q may be expressed as

$$Q_q(\bar{Y},\bar{Z}) = T_q - .5 + \frac{q - \sum_{j=1}^{J}\sum_{i=1}^{n_j} w_j y_{ij}}{\sum_{j=1}^{J}\sum_{i=1}^{n_j} w_j z_{ij}} = T_q - .5 + \frac{q - \sum_{j=1}^{J} \pi_j \bar{Y}_j}{\sum_{j=1}^{J} \pi_j \bar{Z}_j},$$

where q is a fraction indicating the quantile under consideration (e.g. .25 for the 25th quantile), T_q is the integer (or simple) quantile, $y_{ij}=I(\tau_{ij}\leq T_q-1)$, $z_{ij} = I(\tau_{ij}=T_q)$, τ_{ij} is the integer value of tenure for the i,j-th observation, I( ) is again the indicator function, $\bar{Y}_j$ and $\bar{Z}_j$ are the within-group means of y_{ij}

and z_{ij}, $\bar{Y} = \left[\bar{Y}_1,\ldots,\bar{Y}_J\right]$ and $\bar{Z}=\left[\bar{Z}_1,\ldots,\bar{Z}_J\right]$.

Conditioning on T_q, y_{ij} and z_{ij} are Bernoulli random variables with means (μ_{yj} and μ_{zj}) equal to the true shares of the population within groups with tenures that (when rounded to integers) satisfy the expressions in parentheses. Keeping in mind that ($\tau_{ij} \leq T_q$-1), ($\tau_{ij}=T_q$) and ($T_q \geq T_q$+1) are mutually exclusive and exhaustive events, the covariance matrix of y_{ij} and z_{ij} is Σ_j, which has $\mu_{yj}(1-\mu_{yj})$ and $\mu_{zj}(1-\mu_{zj})$ on the diagonal and $-\mu_{yj}\mu_{zj}$ on the off-diagonal. By the Central Limit Theorem, $\bar{Y}_j$ and $\bar{Z}_j$ are asymptotically normal with means μ_{yj} and μ_{zj} and covariance matrix $(1/n_j)\Sigma_j$. If statistics are independent across groups, then the vector $\bar{A} = \left[\bar{X}_1,\bar{Y}_1,\ldots,\bar{X}_J,\bar{Y}_J\right]$ is asymptotically normal with covariance matrix Σ=diag($\Sigma_1,\ldots,\Sigma_J$]. Following the delta method, note that $Q_q(\bar{A}) - Q_q(\mu)$ is asymptotically normal with mean zero and variance $K\Sigma K^T$, where K, which is the vector of partial derivatives of Q_q with respect to $\bar{A}$, is given by

$$K = \left[\pi_1,\ldots,\pi_J\right] \otimes \left[-\frac{1}{\mu_z} \quad \frac{-(q-\mu_y)}{\mu_z^2}\right],$$

and where μ is the vector $[\mu_{y1},\mu_{z1},,,,\mu_{yJ},\mu_{zJ}]$, $\otimes$ is the Kronecker product operator, and $\mu_y = \sum_{j=1}^{J} \pi_j \mu_{yj}$ and $\mu_z = \sum_{j=1}^{J} \pi_j \mu_{zj}$ are the population means of y_{ij} and z_{ij}. The asymptotic variance expression reduces to

$$\frac{1}{\mu_z^2} \sum_{j=1}^{J} \frac{\pi_j^2}{n_j} \left[\mu_{yj}(1-\mu_{yj}) - \frac{2(q-\mu_y)}{\mu_z}\mu_{yj}\mu_{zj} + \left(\frac{q-\mu_y}{\mu_z}\right)^2 \mu_{zj}(1-\mu_{zj})\right]$$

and may be estimated consistently by substituting overall weighted frequencies of y_{ij} and z_{ij} for μ_y and μ_z, and within-group frequencies for μ_{yj} and μ_{zj}.

<u>Standard Errors for Retention Rates.</u> B-year retention rates for jobs in the tenure range $[S_1,S_2]$ and age range $[A_1,A_2]$ in the initial year can be written as follows:

$$R(\bar{U},\bar{V}) = \frac{P_B \sum_{j=1}^{J_B} w_{Bj} \sum_{i=1}^{n_j} v_{ij}}{P_0 \sum_{j=1}^{J_0} w_{0j} \sum_{i=1}^{n_j} u_{ij}} = \frac{P_B \sum_{j=1}^{J_B} w_{Bj} n_j \frac{1}{n_j} \sum_{i=1}^{n_j} v_{ij}}{P_0 \sum_{j=1}^{J_0} w_{0j} n_j \frac{1}{n_j} \sum_{i=1}^{n_j} u_{ij}} = \frac{P_B \sum_{j=1}^{J_B} \pi_{Bj} \bar{V}_j}{P_0 \sum_{j=1}^{J_0} \pi_{0j} \bar{U}_j},$$

where P_0 and P_B are population sizes in years 0 and B, w_{0j} and w_{Bj} are normalized weights as defined above, π_{0j} and π_{Bj} are population shares,

$$u_{ij} = I(S_1 \leq \tau_{ij} \leq S_2, A_1 \leq a_{ij} \leq A_2),$$
$$v_{ij} = I(S_1 + B \leq \tau_{ij} \leq S_2 + B, A_1 + B \leq a_{ij} \leq A_2 + B),$$

I( ) is again the indicator function, τ_{ij} is the value of tenure for the i,jth observation, a_{ij} is the value of age for the i,jth observation, and $\bar{U}_j$ and $\bar{V}_j$ are within-group relative frequencies of individuals satisfying the various sets of qualifications. When the retention rate calculation is done for a subpopulation (i.e. private sector wage employees), the population sizes P_0 and P_B are subpopulation sizes, and the weights w_{0j} and w_{Bj} are normalized by the subpopulation sizes. (Here I am assuming that membership in the subpopulation is independent of job tenure. While not entirely satisfactory, this assumption is consistent with the treatment of subpopulation membership in the main analysis.)

u_{ij} and v_{ij} are Bernoulli random variables with means (μ_{uj} and μ_{vj}) equal to the true shares of the population within groups of individuals with the characteristics described in the inequality expressions. Their variances are $\mu_{uj}(1-\mu_{uj})$ and $\mu_{vj}(1-\mu_{vj})$, and, as they are derived from independent samples, it is reasonable to treat them as independent. By the Central Limit Theorem, $\bar{U}_j$ and $\bar{V}_j$ are asymptotically normal with means μ_{uj} and μ_{vj}, and variances $\mu_{uj}(1-\mu_{uj})/n_j$ and $\mu_{vj}(1-\mu_{vj})/n_j$. If statistics are independent across groups, then the vector $\bar{C} = \left[\bar{U}_1,...,\bar{U}_{J_0},\bar{V}_1,...,\bar{V}_{J_B}\right]$ is asymptotically normal with mean vector $\mu^* = \left[\mu_{u1},...,\mu_{uJ_0},\mu_{v1},...,\mu_{vJ_B}\right]$ and covariance matrix $\Omega = \text{diag}\left[\mu_{u1}(1-\mu_{u1})/n_1,...,\mu_{uJ_0}(1-\mu_{uJ_0})/n_{J_0},\mu_{v1}(1-\mu_{v1})/n_1,...,\mu_{vJ_B}(1-\mu_{vJ_B})/n_{J_B}\right]$.

Following the delta method, note that $R(\bar{C}) - R(\mu*)$ is asymptotically normal with covariance $L\Omega L^T$, where L, which is the vector of partial derivatives of R with respect to $\bar{C}$ may be expressed

$$L = \frac{P_B}{P_0\mu_u}\left[\frac{-\mu_v}{\mu_u}\pi_{01}\cdots\frac{-\mu_v}{\mu_u}\pi_{0J_0}\ \pi_{B1}\cdots\pi_{BJ_B}\right]$$

and $\mu_u = \sum_{j=1}^{J_0} \pi_j \mu_{uj}$ and $\mu_v = \sum_{j=1}^{J_B} \pi_j \mu_{vj}$ are the population means of u_{ij} and v_{ij}. This reduces to

$$(\frac{P_B}{P_0\mu_u})^2\left[(\frac{\mu_v}{\mu_u})^2\sum_{j=1}^{J_0}(\pi_{0j})^2\mu_{uj}(1-\mu_{uj})/n_j+\sum_{j=1}^{J_B}(\pi_{Bj})^2\mu_{vj}(1-\mu_{vj})/n_j\right]$$

and may be estimated consistently by substituting overall weighted frequencies of u_{ij} and v_{ij} for μ_u and μ_v, and within-group frequencies for μ_{uj} and μ_{vj}.

Table 1
Descriptive Statistics [1]
Males Aged 15 to 59

	Colombia 1988	United States 1987
Population Represented (Millions)	3.25	54.11
Number of Observations	28,528	30,045
Means:		
Age in years	31.5	34.1
Schooling in Years	8.2	12.7
Percent of Population 15-59:		
Aged 15-20	20	15
With some schooling	98	97
With secondary complete	32	77
With more than secondary	16	43
Employed	76	75
Unemployed	7	3
Out of Labor Force	16	22
Percent of Employed in:		
Private Wage Jobs	57	75
Public Sector	10	13
Self Employment	33	12
White Collar Occupations	21	38
Manufacturing	24	23
Food, Textiles	9	3
Construction	9	11
Commerce	25	20
Transport, Communication	9	8
Finance, Real Estate	7	13
Services	22	21
Personal Services	11	2
Other	2	2

1. Calculations employ population weights as supplied by DANE and the CPS.

Table 2
Current Tenure Statistics [1]
Males Private Sector Wage Employees

	Colombia 1988	United States 1987
Number of Observations	11,316	15,686
Current Tenure in Years:		
Mean	4.27	6.62
90th Quantile	12.47	18.49
75th Quantile	5.34	9.59
50th Quantile	1.87	3.51
25th quantile	0.39	1.00
10th Quantile	0.15	0.24
Percent of Workers with Years of Tenure:		
Over 20	4	7
11-20	8	14
6-10	12	18
3-5	18	20
1-2	25	21
Zero	32	21

1. Calculations in this and all subsequent tables employ weights as supplied by DANE and the CPS, adjusted following the discussion in the appendix. Tenure quantiles here and in all subsequent tables are interpolated quantiles as described in the text. Standard errors on the interpolated quantiles rise from .002 (.003) to .211 (.152) as one moves from the 10th to the 90th quantile for Colombia (the United States).

Table 3
Interpolated Median Job Tenure by Age
Male Private Sector Wage Employees

	Colombia 1988	United States 1987
Aged 15-20	0.44	0.44
Aged 21-30	1.50	2.09
Aged 31-50	3.25	4.67
Aged 41-50	5.56	9.10
Aged 51-59	8.50	13.95

1. Standard errors for the Colombian(U.S.) interpolated medians rise with age bracket, from .009 (.013) for ages 15-20 to .477 (.370) for ages 41-50 and then rise more abruptly to 2.70 (.470) for ages 51-59.

Table 4
Four-Year Job Retention Probabilities by Initial Tenure [1]
Male Private Sector Wage Employees Aged 15-59 in Initial Year

	Colombia		United States
Initial Tenure	1984-1988	1988-92	1987-91
0	.19	.18	.31
1-2	.39	.39	.52
3-5	.36	.40	.39
6-10	.63	.80	.65
11-20	.94	.90	.80
>20	.96	.59	.65

1. In general, estimated standard errors rise with initial tenure. For Colombian in 1984-88 (1988-92), they rise from .008 (.008) at tenure 0 to .068 (.047) at tenure over 20. For the United States comparable errors rise from.010 to .027.

Table 5
Four-Year Job Retention Probabilities by Initial Tenure and Initial Age [1]
Male Private Sector Wage Employees

Initial Age	Initial Tenure in Years			
	0	1-2	3-5	6-10
Colombia 1984-88				
Ages 15-20	.18	.39	.52	.
Ages 21-30	.19	.35	.35	.72
Ages 31-40	.19	.47	.35	.61
Ages 41-50	.20	.42	.39	.60
Ages 51-59	.15	.42	.15	.37
Columbia 1988-92				
Ages 15-20	.15	.36	.37	.
Ages 21-30	.19	.37	.37	.80
Ages 31-40	.21	.46	.47	.89
Ages 41-50	.16	.43	.46	.79
Ages 51-59	.09	.35	.20	.45
United States 1987-91				
Ages 15-20	.25	.46	.37	.
Ages 21-30	.30	.52	.38	.69
Ages 31-40	.35	.57	.43	.69
Ages 41-50	.37	.55	.43	.65
Ages 51-59	.24	.35	.29	.40

1. In general estimated standard errors are smallest, on the order of .01 (.02) in Colombia (the United States), at low tenures for young ages. The highest standard errors for retention rates in this table are on the order of .08.

Table 6
Four-Year Job Retention Probabilities[1]
Colombian Male Private Sector Wage Employees

Initial Tenure	1984-88	1988-92	1992-96
0	.19	.18	.20
1-2	.39	.39	.34
3-5	.36	.40	.34
6-10	.63	.80	.51
11-20	.94	.90	.72
> 20	.96	.59	.81

[1] In general, estimated standard errors rise with initial tenure, from the range of .008-.010 at initial tenure of zero to the range of .04-.07 at initial tenure over 20.

Table 7
Two-Year Job Retention Probabilities [1]
Colombian Male Private Sector Wage Employees

Initial Tenure	1984-86	1986-88	1992-94	1994-96
0	.40	.41	.43	.44
1-2	.63	.59	.57	.53
3-5	.63	.57	.57	.56
6-10	.86	.71	.71	.71
11-20	1.01	.86	.82	.86
>20	.91	1.05	.97	.81

1. In general standard errors rise with initial tenure and are fairly stable over time for the same initial tenure. They rise from roughly .01 at initial tenure of zero to .07 at initial tenures over 20 .

Table 8
Weighted Least Squares Regression of Interpolated Medians on Cell Characteristics[1]
Male Private Sector Wage Employees

Cells Defined By:	Age, Country, Industry, Occupation	Age, Country, Industry, Occupation, Size
Ages 15-20	0.43 (0.03)	0.46 (0.02)
Ages 21-30	1.49 (0.09)	1.36 (0.11)
Ages 31-40	2.57 (0.25)	1.34 (0.55)
Ages 41-50	3.39 (0.99)	0.25 (1.64)
Ages 51-59	5.47 (0.99)	2.05 (1.80)
US=United States		
US*Ages 15-20	-0.00 (0.07)	-0.04 (0.03)
US*Ages 21-30	0.58 (0.15)	0.58 (0.12)
US*Ages 31-40	1.39 (0.69)	0.95 (0.72)
US*Ages 41-50	2.61 (1.11)	2.26 (1.27)
US*Ages 51-59	2.24 (1.86)	1.31 (1.97)
Industry-Age Interactions	Yes	Yes
W=White Collar		
W*Ages 15-20	0.08 (0.15)	-0.02 (0.11)
W*Ages 21-30	0.36 (0.18)	0.34 (0.16)
W*Ages 31-40	1.15 (0.70)	0.78 (0.74)
W*Ages 41-50	2.28 (1.02)	1.71 (1.20)
W*Ages 51-59	3.80 (2.60)	2.24 (2.19)
L=Large		
L*Ages 15-20		-0.03 (0.02)
L*Ages 21-30		0.25 (0.10)
L*Ages 31-40		2.58 (0.52)
L*Ages 41-50		5.05 (1.54)
L*Ages 51-59		8.35 (2.16)
R-Squared	.96	.94
Number of Cells	125	166
P-Value for test of No U.S. Differences	.0003	.0001

[1] The regressions employ 1988 data from both Colombia and the United States. They are weighted by the number of observations in the cell. Heteroskedasiticity-consistent standard errors are reported in parentheses. Cells with fewer than 20 observations are excluded. There are 8 industry categories, 2 occupation categories, 2 size categories and 2 education categories. The excluded industry is commerce. The intercept has been suppressed so that all age categories may be included.

Table 9
Weighted Least Squares Regressions of Four-Year Job Retention Probabilities on Cell Characteristics[1]
Male Private Sector Wage Employees

Cells Defined By:	**Tenure, Country, Industry, Occupation**
Tenure 0	0.15 (0.02)
Tenure 1-2	0.34 (0.05)
Tenure 3-5	0.33 (0.03)
Tenure 6-10	0.68 (0.04)
Tenure 11-20	0.78 (0.60)
Tenure > 20	0.60 (0.07)
US=United States	
US*Tenure 0	0.11 (0.02)
US*Tenure 1-2	0.13 (0.04)
US*Tenure 3-5	-0.02 (0.04)
US*Tenure 6-10	-0.14 (0.04)
US*Tenure 11-20	-0.10 (0.05)
US*Tenure > 20	0.01 (0.05)
Industry-Tenure Interactions	Yes
W=White Collar	
W*Tenure 0	0.17 (0.03)
W*Tenure 1-2	0.06 (0.05)
W*Tenure 3-5	0.09 (0.04)
W*Tenure 6-10	0.06 (0.07)
W*Tenure 11-20	0.10 (0.06)
W*Tenure > 20	0.11 (0.04)
R-Squared	.98
Number of Cells	162
P-Value of test of No U.S. Differences	.0000

[1] Colombian (U.S.) retention rates are for the period 1988-92 (1987-91). The regressions are weighted by the number of observations in the cell. Heteroskedasticity-consistent standard errors are reported in parentheses. Cells with fewer than 20 observations are excluded. There are 8 industry categories, 2 occupational categories and 2 education categories. The excluded industry is commerce. The intercept has been suppressed so that all initial tenure categories may be included.

Figure 1
Percentage of Male Workers in Private Sector Wage Employment
by Age, Year and Country