Pergamon
www.elsevier.com/locate/worlddev
World Development Vol. 29, No. 4, pp. 725–745, 2001

Printed in Great Britain
0305-750X/01/$ - see front matter
PII: S0305-750X(00)00117-0

Patterns and Sources of Changing Wage Inequality in Chile and Costa Rica During Structural Adjustment

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Summary. — During their respective periods of structural adjustment, inequality increased more rapidly in Chile than in Costa Rica. Using a new technique which measures the effects of changes in the quantities and prices of individual dimension of human capital on overall wage inequality, we identify changes in wage premiums associated with more education as an important cause of the different inequality outcomes. We present evidence that the different education price effects were due to different rates of growth in the demand for more-educated workers. Furthermore, inequality increased in Chile despite a large equalizing education quantity effect.

Key words — income inequality, structural adjustment, trade liberalization, education, wages, Latin America

1. INTRODUCTION

This paper documents, compares and analyzes the outcomes and causes of changes in the distribution of real hourly wages in Chile over 1974–90 and Costa Rica over 1987–95. Chile and Costa Rica are similar in many respects. Both are middle-income developing countries, both have a high proportion of salaried employees and a highly educated work force, and both have largely resource-based exports with moderate industrial and service sectors. [1] In addition, both countries have experienced rapid educational expansion and have relatively recently undergone major trade liberalization and other structural adjustment reforms. Despite these similarities, during recent periods of structural adjustment inequality deteriorated rapidly in Chile, while improving or remaining constant in Costa Rica.

Robbins (1994) and Robbins and Gindling (1999) document an increase in the relative wages of more-educated workers that coincided with structural adjustment in both Chile and Costa Rica, with a much more rapid increase in Chile compared to Costa Rica. [2] An implication of this earlier work is that different rates of change in relative wages in Chile and Costa Rica might explain the different inequality outcomes. But while rising relative wages may have contributed to the different inequality outcomes, their effect may have been small compared to other factors. For example, the differences in inequality outcomes, especially the widening wage distribution in Chile, could also have arisen from the disequalizing effects of education expansion on the composition of the work force. Educational expansion may affect the wage distribution in two ways, which Knight and Sabot (1983) call the "composition" effect and the "wage-compression" effect. The composition effect occurs because of changes

* We would like to thank Albert Berry, Chinhui Juhn, Juan Diego Trejos and three anonymous referees for comments and help. All errors are our own. Final revision accepted: 9 September 2000.

in the distribution of education in the working population (holding the wages of workers at different education levels constant). *Ex ante*, the impact of the composition effect on inequality is indeterminate. The wage-compression effect occurs because the increase in the supply of more-educated workers pushes the wages of more-educated workers down relative to the wages of less-educated workers (holding the distribution of education constant). The wage-compression effect has an unambiguously equalizing effect on the distribution of wages. [3]

In this paper we measure the degree to which changes in the distribution and wages of more-educated workers, as opposed to changes in the distribution and wages of other wage-determining characteristics such as experience, caused the changes in wage inequality in Chile and Costa Rica. To document and analyze changes in wage distributions we adopt a modification of the methodology presented in Juhn, Murphy, and Pierce (1993). The Juhn *et al.* technique decomposes the change in inequality into three parts: a part due to changes in the distribution of human capital, a part due to changes in the wage premiums (prices of labor) associated with this human capital, and a part due to changes in unobserved quantities and prices. We modify the Juhn *et al.* (1993) technique in a way that permits us to estimate separately the quantity and price effects of individual dimensions of observable human capital (for example, education). Decomposing wage inequality into price and quantity effects for individual dimensions of human capital is not something that has been done in previously published research, although the technique we use is similar to techniques presented in recent working papers by Bourguignon, Fournier, and Gurgand (1998) and Fields (1998). [4]

The rest of the paper is organized as follows: Section 2 presents the data and patterns of changes in inequality. Section 3 presents the technique for decomposing changes in wage inequality. Section 4 presents the empirical results. Section 5 concludes. We identify changes in wage premiums associated with more education as an important cause of the different inequality outcomes. We present evidence that the different education price effects were due to different rates of growth in the demand for more-educated workers. Furthermore, inequality increased in Chile despite a large equalizing education quantity effect.

2. DATA AND PATTERNS OF CHANGES IN INEQUALITY

The data used are from yearly household surveys. For Chile the data are from the University of Chile Household Surveys of Greater Santiago which have been conducted in June since 1957. The Greater Santiago area constitutes roughly 40% of Chile's population and over half of its output. The average number of individuals sampled is approximately 10,000, of which the active labor force averages roughly 4,000. For Costa Rica the data are from similar household surveys, the Household Survey for Multiply Purposes (called the Household Surveys of Employment and Unemployment prior to 1987). The Costa Rican household surveys have been conducted countrywide each July since 1976 by the Directorate of Statistics and Census. [5] Approximately 0.5–1% of the population is sampled each year. The Chilean and Costa Rica surveys ask similar questions and report similar information on earnings, hours worked, education level, employment, sex, age and other labor market and demographic variables.

In this paper we focus on comparing periods of structural adjustment in Chile and Costa Rica. The timing of the reforms was different in the two countries: structural adjustment began in the early 1970s in Chile and in the mid-1980s in Costa Rica; specifically, we compare Chile during 1974–90 and Costa Rica during 1987–95. [6]

Figure 1 presents changes in the variance of the log of real wages in Costa Rica and Chile during the years of structural adjustment. [7] During structural adjustment in Chile the variance increased, while with structural adjustment in Costa Rica the variance changes very little. [8]

Figure 2 presents a more detailed disaggregation of the change in inequality—the log ratio of hourly wages of workers at the 90th percentile in the log wage distribution divided by the hourly wages of workers at the 10th percentile (the 90/10 ratio), the 90/50 ratio, and the 50/10 ratio. Table 1 presents the changes in the log of real hourly wages at the 90th, 50th and 10th percentiles. In Chile, the increase in inequality during the period of adjustment occurs because of more rapid increases in the income of the

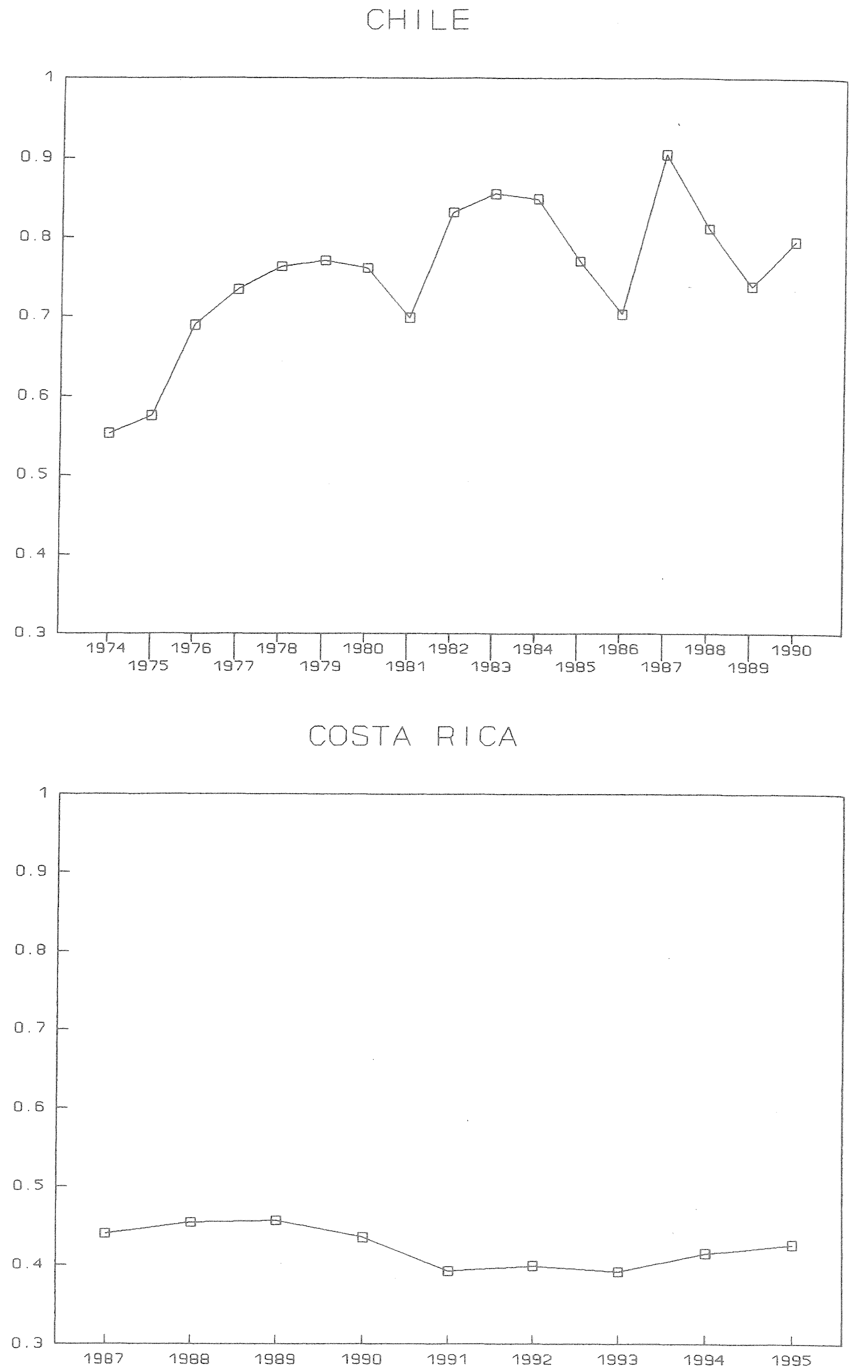

Figure 1. *Variance of the log of real hourly wages for male salaried employees: Chile (1974–90) and Costa Rica (1987–95).*

90th percentile versus the middle and the bottom (the 50th and 10th percentiles)—the log of real hourly wages of workers at the 90th percentile increases by more than six times the increase of the log of hourly wages of workers at the 50th and 10th percentiles. As a result, the 90/10 and 90/50 ratios increase while the 50/10 ratio changes little during 1974–90 (and actually falls during 1975–90). On the other hand, in Costa Rica, the middle-income workers fare better compared to workers at the 90th percentile—the increase in the log of real wages of workers

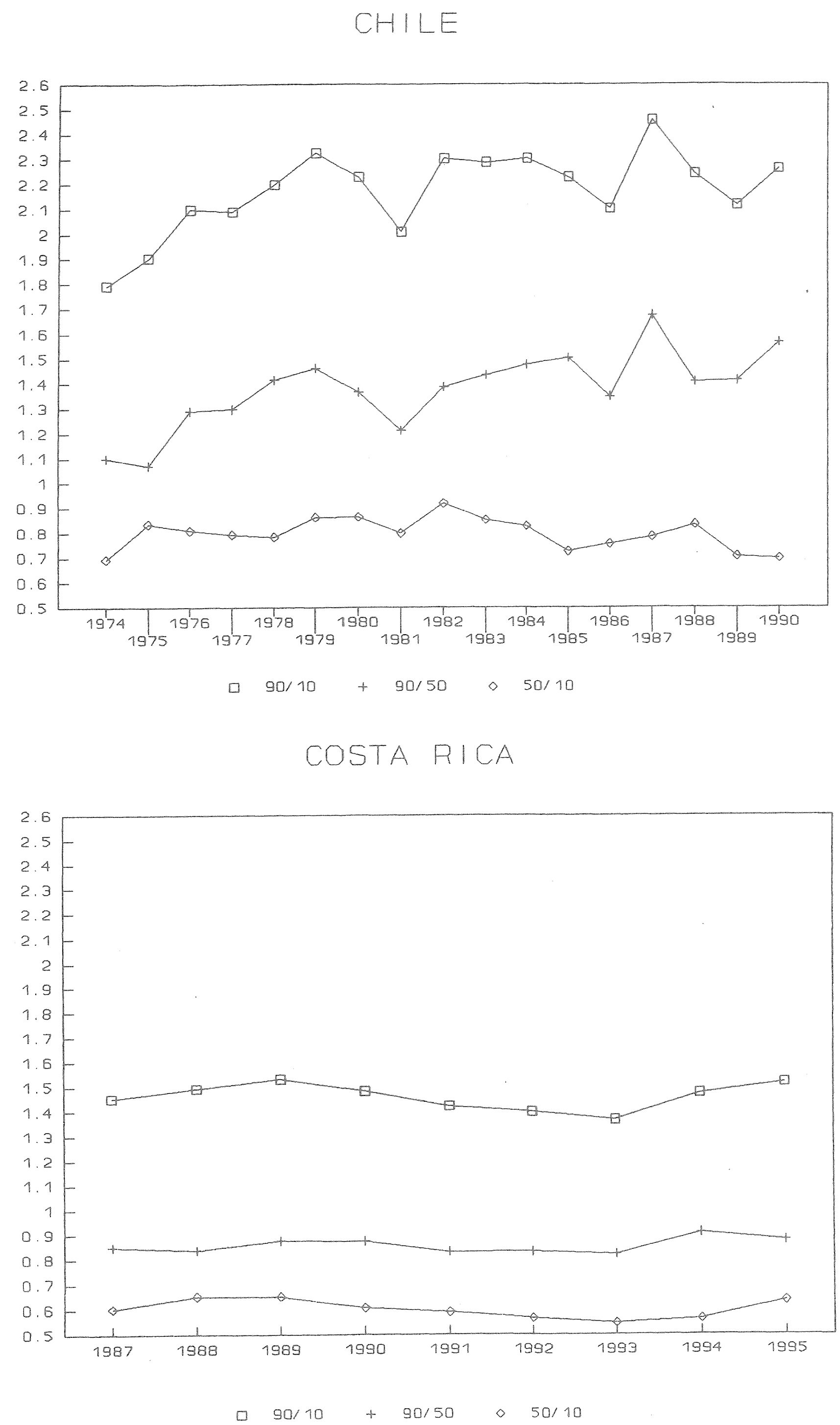

Figure 2. *Log ratios of the real hourly wages for male salaried empolyees at the 90th, 50th, and 10th centiles: Chile (1974–90) and Costa Rica (1987–95).*

at the 90th percentile in Costa Rica is less than twice that of workers at the 50th or 10th percentile. As a result, the 90/10 and 90/50 ratios increase more in Chile than in Costa Rica, while the 50/10 ratio increases less in Chile than in Costa Rica. [9]

Table 1. *Change in the log of real hourly wages for male salaried workers of the 90th, 50th and 10th percentiles in Chile (1974–90) and Costa Rica (1987–95)*

	Chile, 1974–90		Costa Rica, 1987–95	
	Overall change	Implicit annual change	Overall change	Implicit annual change
90th percentile	0.558	0.035	0.114	0.014
50th percential	0.091	0.006	0.085	0.011
10th percential	0.087	0.005	0.051	0.006

3. TECHNIQUE FOR DECOMPOSING CHANGES IN EARNINGS INEQUALITY

We use a modification of a technique developed by Juhn *et al.* (1993). The Juhn *et al.* technique decomposes the change in inequality into three parts: a part due to changes in the distribution of human capital (quantities), a part due to changes in the wage premiums (prices) associated with this human capital, and a part due to changes in unobserved or unmeasured quantities and prices. Their technique is based on the estimation of the wage equations

$$Y_{it} = X_{it}B_t + u_{it}, \tag{1}$$

where Y_{it} represent the (log) of real wages for individual i in year t, X_{it} is a vector of individual human capital characteristics (including education and experience) for individual i in year t, the elements of the vector B_t can be interpreted as the wage premium, or price, for each human capital characteristics in year t, and u_{it} is the component of wages accounted for by factors we cannot measure or observe.

The Juhn *et al.* technique begins by considering an average (over-all-years) distribution of wages, Y_i, where

$$Y_i = \mathbf{X}_i\mathbf{B} + \mathbf{u}_i, \tag{2}$$

where $\mathbf{X}_i$ is a vector of the average (over-all-years) human capital characteristics for individual i, $\mathbf{u}_i$ are average (over-all-years) errors, and $\mathbf{B}$ is a vector of the average (over-all-years) wage equation coefficients.[10] Beginning with the average wage distribution represented by Eqn. (2), they then: first, replace $\mathbf{X}_i$ with X_{it} for each individual i (to construct a distribution of wages where the human capital characteristics are those in year t but prices and the distribution of errors are the over-all-years averages); second, replace $\mathbf{B}$ with B_t (to construct a distribution of wages where human capital characteristics and prices are those in year t but the distribution of errors is the over-all-years average); finally replace $\mathbf{u}_i$ with u_{it} (resulting in the actual distribution of wages in time t—represented by Eqn. (1)).

Our modification of the Juhn *et al.* technique allows us to estimate quantity and price effects separately for different types of human capital. To do this, we maintain the sequential nature of the Juhn *et al.* technique; starting with the average distribution represented by Eqn. (2) and ending with the actual distribution represented by Eqn. (1). Starting with the average distribution, we first sequentially replace each component of the vector of average human capital characteristics, $\mathbf{X}_i$, with the actual characteristic in time t for each individual i, X_{it}. We then, in turn, change the price for each human capital characteristic from the average price, $\mathbf{B}$, to the price in time t, B_t. Finally, we replace $\mathbf{u}_i$ with u_{it}.

More formally, we estimate a wage equation with right-hand-side variables representing two types of human capital: education and experience.[11] We first construct the distribution of wages for each year t where the distribution of education is that in year t, but the distribution of experience, errors and the prices are the over-all-year averages,

$$Y_{it}^1 = X_{it}^a\mathbf{B}^a + \mathbf{X}_i^b\mathbf{B}^b + \mathbf{u}_i, \tag{3}$$

where "a" denotes the education variables and "b" denotes the experience variables. The average (over-all-years) prices, $\mathbf{B}^a$ and $\mathbf{B}^b$, were estimated with a pooled regression using data from all years. Following the procedure outlined in Juhn *et al.* (1993), we compute $\mathbf{u}^i$, the average residual for each percentile group in the distribution of residuals, based on the workers actual percentile in each year's residual distribution and the average cumulative distribution over the full sample (for all years). $\mathbf{X}_i^b$ is a vector of experience variables based on the average (over-all-years) experience for individual i's percentile in the distribution of experience. Using a procedure similar to that used to calculate the average residual distribution, $\mathbf{X}_i^b$

was computed based on the workers actual percentile in each year's experience distribution and the average cumulative distribution of experience in the full sample (for all years).

Next, we sequentially change each component of the wage equation and calculate the distribution of wages in each case. [12] Specifically, we calculate

$$Y_{it}^2 = X_{it}^a \mathbf{B}^a + X_{it}^b \mathbf{B}^b + \mathbf{u}_i, \tag{4}$$

$$Y_{it}^3 = X_{it}^a B_t^a + X_{it}^b \mathbf{B}^b + \mathbf{u}_i, \tag{5}$$

$$Y_{it}^4 = X_{it}^a B_t^a + X_{it}^b B_t^b + \mathbf{u}_\mathbf{i}, \tag{6}$$

$$Y_{it} = X_{it}^a B_t^a + X_{it}^b B_t^b + u_{it}. \tag{7}$$

Changes over time in inequality in the distribution of Y_{it}^1 are an estimate of changes in wage inequality allowing the distribution of education to change while holding constant: the distribution of experience, the prices of experience and education, and the distribution of unobservables. We call this the contribution of "education quantities" to changes in the distribution of wages. The education quantity effect is an estimate of the composition effect of educational expansion.

Changes over time in inequality in the distribution of Y_{it}^2 are an estimate of changes in wage inequality allowing the distribution of both types of human capital to change while holding constant: education and experience prices, and unobservables. Any additional changes in inequality in the distribution of Y_{it}^2 beyond changes in inequality in the distribution of Y_{it}^1 capture changes in wage inequality due to changes in the distribution of experience holding all else constant. [13] We call this the contribution of "other quantities" to changes in the distribution of wages. [14]

Any additional changes in inequality in the distribution of Y_{it}^3 beyond changes in inequality in the distribution of Y_{it}^2 capture changes in wage inequality due to changes in education prices, holding all else constant. We call this the contribution of "education prices" to changes in the distribution of wages.

If the relative demand for workers with different types of education is constant, then the education price effect is a precise estimate of the wage-compression effect of educational expansion. But if both demand and supply are changing, then the measured education price effect reflects the impact of both educational expansion and of changes in demand.

Any additional changes in inequality in the distribution of Y_{it}^4 beyond changes in inequality in the distribution of Y_{it}^3 capture changes in wage inequality due to changes in experience prices holding all else constant. We call this the contribution of "other prices" to changes in the distribution of wages. [15]

Finally, additional changes in inequality between Y_{it} and Y_{it}^4 can be attributed to changes in the distribution of the error terms (quantities and prices that we do not observe). We call this the "contribution of unobservables." Juhn *et al.* interpret changes in the unobservables as changes in unobservable skill quantities and prices.

To control for possible data problems with using data from women and self-employed workers, following Juhn *et al.* (1993) we estimate the decompositions using data from only male salaried employees who report their wage, education and age. We decompose the variance and the 90/10, 90/50 and 50/10 ratios.

4. WAGE INEQUALITY DECOMPOSITIONS RESULTS

Table 2 presents the wage decomposition for the adjustment periods in Chile and Costa Rica. [16] For comparison purposes the most important columns in this table are the "implicit annual change" columns, and we will refer to these columns in the text. A negative number means that the contribution (or effect) is equalizing; a positive number means that the contribution is disequalizing. Overall, inequality rose much faster in Chile than in Costa Rica. In Chile, using all inequality measures, inequality of hourly wages grew rapidly coincident with structural adjustment. In Costa Rica, depending on the measure used, inequality of hourly wages either fell (using the variance) or grew slightly (using centile ratios). While the 90/10 and 90/50 ratios increase in both Costa Rica and Chile, the annual average increases are larger in Chile than Costa Rica. The 50/10 ratio also rises in both countries, but the increase is larger in Costa Rica.

The most robust results that one can derive from Table 1 are where the signs of the effects are the same for all four measures of inequality. Using this criteria, we conclude that in Chile changes in education prices, experience prices, and unobservables had a disequalizing effect on the wage distribution, while changes in experience quantities were

Table 2. *Wage inequality decompositions during structural adjustment: in Chile (1974–90) and Costa Rica (1987–95)*[a]

	Chile, 1974–90		Costa Rica, 1987–85	
	Overall change	Implicit annual change	Overall change	Implicit annual change
Variance				
Total	0.242	0.015	−0.014	−0.002
Total observable quantities	−0.076	−0.005	0.021	0.003
Education quantities	−0.073	−0.005	0.038	0.005
Other quantities	−0.003	−0.000	−0.017	−0.002
Total observable prices	0.194	0.012	−0.017	−0.002
Education prices	0.149	0.009	−0.031	−0.004
Other prices	0.045	0.003	0.014	0.002
Unobservables	0.125	0.008	−0.018	−0.002
90/10 Ratio				
Total	0.470	0.029	0.063	0.008
Total observable quantities	−0.118	−0.007	0.077	0.010
Education quantities	−0.096	−0.006	0.073	0.009
Other quantities	−0.022	−0.001	0.003	0.000
Total observable prices	0.321	0.020	−0.026	−0.003
Education prices	0.239	0.015	−0.064	−0.008
Other prices	0.083	0.005	0.038	0.005
Unobservables	0.266	0.017	0.012	0.002
90/50 Ratio				
Total	0.466	0.029	0.029	0.004
Total observable quantities	0.014	0.001	0.063	0.008
Education quantities	0.026	0.002	0.056	0.007
Other quantities	−0.013	−0.001	0.007	0.001
Total observable prices	0.273	0.017	−0.026	−0.003
Education prices	0.201	0.013	−0.046	−0.006
Other prices	0.071	0.004	0.020	0.002
Unobservables	0.180	0.011	−0.007	−0.001
50/10 Ratio				
Total	0.004	0.000	0.033	0.004
Total observable quantities	−0.132	−0.008	0.014	0.002
Education quantities	−0.122	−0.008	0.018	0.002
Other quantities	−0.009	−0.001	−0.004	−0.000
Total observable prices	0.049	0.003	−0.000	−0.000
Education prices	0.038	0.002	−0.018	−0.002
Other prices	0.011	0.001	0.018	0.002
Unobservables	0.086	0.005	0.020	0.002

[a] Changes in the contribution of observable education and observable experience quantities and prices, and unobservables, to the total change in wage inequality.

equalizing. In Costa Rica, changing education prices were equalizing while changing education quantities were disequalizing.[17]

The criteria that the signs be the same for all four measures of inequality has two important limitations: (a) because the units used in each inequality measure are different, it is difficult to make conclusion about the average (across all inequality measures) relative magnitude of each effect—for example, in Chile we cannot tell whether the education price effect was more or less important than the unobservables effect in causing the increase in inequality; (b) when the signs of the effects differ depending on the measure of inequality, we would like to be able to conclude something about the average direction of the effect. To address these limitations, we present two summary measures that, while less robust that the same-sign criteria, allow us to make conclusions about the average relative magnitude of each effect.

In order to make summary comparisons about the relative magnitude of the different effects, the first summary measure reported is the mean rank ordering of the basic effects. This measure must employ the condition that the signs be alike in all inequality measures (where signs of effects differ over inequality measures the value is reported as missing, or "."). To construct this measure we first calculate the ranking of the absolute value of the magnitude of each effect within each inequality measure; assigning a rank of one to the effect with the largest absolute value, two to the next largest, and so on. Then we calculate the average rank of each of these effects across inequality measures. This summary measure is intended to be a conservative measure of common tendencies over inequality measures that only compare the order of within-measure effects. This summary measure has the advantage of not making cardinal comparisons of effects across inequality measures (which are not strictly cardinally comparable). The advantage of requiring signs be uniform for all inequality measures, and using the mean rank orderings subject to that criterion, is robustness.

While the average rank orderings permit ordinal comparisons over groups, they do not allow cardinal comparisons of the importance of effects across groups, and they eliminate comparisons of effects that do not have like signs over all groups. Thus, at the sacrifice of some robustness, we also present a second summary measure that is the average of normalized effects over inequality measures. The normalized effect for each measure of inequality is calculated by dividing individual effects by the sum of the absolute values of all effects (education quantity effect plus other quantity effect plus education price effect plus other price effect plus unobservables effect). We then report the mean of these normalized effects over all four inequality measures.

These two summary measures are presented in Table 3. The two measures reported are consistent with each other: the ranking of effects using either the mean rank ordering or the mean of normalized effects is nearly identical. When we refer to Table 3 in the discussion below, we will focus on the mean normalized effects columns.

(a) *Chile*

The summary average measures reported in Table 3 show that the disequalizing education price and unobservables effects were quantitatively the most important factors causing the deterioration in wage inequality in Chile. The next largest effect was the equalizing education quantities effect, followed by the disequalizing other price effect, and the equalizing other quantities effect (in that order).

The results of the decompositions lead us to conclude that the large widening in the distribution of wages in Chile was in large part due to the education price effect. The disequalizing effect of education prices implies that the wages of more-educated workers rose relative to the wages of less-educated workers. Education prices increased despite rapid increases in the supply of more-educated workers (Table 4), which, in theory, should have contributed to a drop in education prices. The fact that education prices rose in the face of increases in supply indicates that there must have been an increase in demand for more-educated workers. The consistently disequalizing effects of education prices, experience prices, and unobservables in Chile can be interpreted as evidence of an increase in the demand for more-skilled workers in general that accompanied structural adjustment in Chile. Because the trend of the unobservables effect was similar to the observable skill–price effect, Juhn *et al.* (1993) interpret the disequalizing effect of unobservables in the United States as being due to the disequ-

Table 3. *Summary measures of the inequality decompositions*[a]

Effects	Chile, 1974–90		Costa Rica, 1987–95	
	Mean rank ordering	Mean normalized effects	Mean rank ordering	Mean normalized effects
Education quantities	"."	−0.92	1.25	1.69
Other quantities	4	−0.1	"."	−0.16
Education prices	1.5	1.59	2.25	−1.46
Other prices	3	0.52	2.5	0.87
Unobservables	1.5	1.74	"."	0.14

[a] Mean rank ordering is the order of normalized effects where those effects have the same sign across groups. The *j*th normalized effects is: effect(j)/ $\sum_i$(abs(Effect$_i$)). The average normalized effects are simply the average effects over outcome measure groups, without requiring uniformity of signs across groups.

Table 4. *Changes in the mean and variance in years of education*

	Chile, 1974–90		Costa Rica, 1987–95	
	Overall change	Implicit annual change	Overall change	Implicit annual change
Change in the variance in years of education	−3.01	−0.19	0.45	0.06
Change in the mean years of education	2.83	0.18	0.48	0.06

alizing effects of unobservable skill prices. The decompositions we present are consistent with a similar interpretation of the unobservable effect in Chile.

Ex ante, the disequalizing effect of education and other human capital prices in Chile may have been small relative to other factors. For example, the composition effect of educational expansion could have been strongly equalizing. As we can see from Table 3, however, using the mean normalized effects, the average education-quantity effect in Chile was large and equalizing. The result that the measured composition effect of educational expansion was equalizing, along with an *ex ante* theoretical prediction that the wage-compression effect of educational expansion is always equalizing, implies that the overall effects of educational expansion were strongly equalizing in Chile. By implication, the overall deterioration of the distribution of wages in Chile was clearly due to forces other than educational expansion.

In summary, our results indicate that the deterioration in distribution of wages in Chile was driven by education price and unobservables effects. The fact that average educational levels and relative supply were growing rapidly while the overall education price effect was disequalizing implies that during structural adjustment relative demand shifts were strongly skill-biased, and dominated the growth in the supply of education's effects upon wages. Moreover, we have learned that the education price effect was so large that it, along with unobservables effects, dominated the wage-compression and very large equalizing composition effects of educational expansion.

For comparison purposes, Table 5 presents the decompositions for the pre-adjustment period. The education price and quantity effects were very different in the pre-adjustment periods from what they were under structural adjustment. Specifically, education price effects in the pre-adjustment period were equalizing, while education quantity effects in the pre-adjustment period were disequalizing—the opposite of their impacts in the adjustment period. This is additional evidence that the increase in the relative demand for more-educated workers in Chile that coincided with structural adjustment did not simply represent a continuation of underlying trends, but was due to something that happened around 1974, the beginning of the structural adjustment program in Chile.

(b) *Costa Rica*

Compared to Chile, inequality changed little in Costa Rica after structural adjustment. Quantitatively, the most important measured effects were an equalizing education price effect and a disequalizing education quantities effect. The overall effect of educational expansion on the distribution of wages in Costa Rica was roughly neutral. After the education quantity and price effects, the quantitatively next most important effects were (in order) disequalizing other price effects, equalizing other quantities effects and disequalizing unobservables effects.

The disequalizing education quantities effect was on average slightly higher than the measured equalizing education price effect. We suspect, however, that the wage-compression effect of educational expansion in Costa Rica was likely somewhat larger than the measured education price effects. This is for several reasons: first, our prior work found that the relative demand for education grew slightly in Costa Rica in the period of structural adjustment; second, the mean level of education, reported in Table 4, grew over this period. Thus, the measured education price effect likely reflects the equalizing effects from the rising supply of education and small but disequalizing effects of rising relative demand.

Both the signs and the magnitudes of the education quantity effect are similar in the pre-adjustment and the adjustment period in Costa Rica, suggesting that the disequalizing education quantity effect during the period of structural adjustment may have been a continuation of an underlying trend (see Tables 2 and 5).[18] On the other hand, while the signs of the education price effect in Costa Rica were the same

Table 5. *Wage inequality decompositions in the pre-adjustment period: Chile (1957–74) and Costa Rica (1976–85)*[a]

	Chile, 1957–74		Costa Rica, 1976–85	
	Overall change	Implicit annual change	Overall change	Implicit annual change
Variance				
Total	0.004	0.000	−0.103	−0.011
Total observable quantities	0.014	0.001	−0.081	−0.089
Education quantities	0.045	0.003	0.062	0.007
Other quantities	−0.031	−0.002	−0.142	−0.016
Total observable prices	−0.069	0.005	0.009	0.001
Education prices	−0.095	−0.007	−0.136	−0.015
Other prices	0.026	0.002	0.144	0.016
Unobservables	0.059	0.004	−0.031	−0.003
90/10 Ratio				
Total	−0.018	−0.001	−0.288	−0.032
Total observable quantities	0.165	0.013	0.032	0.004
Education quantities	0.191	0.015	0.061	0.007
Other quantities	−0.027	−0.002	−0.029	−0.003
Total observable prices	−0.168	−0.013	−0.230	−0.025
Education prices	−0.150	−0.012	−0.250	−0.028
Other prices	−0.018	−0.001	0.021	0.002
Unobservables	−0.015	−0.001	−0.090	−0.010
90/50 Ratio				
Total	0.118	0.009	−0.237	−0.026
Total observable quantities	0.182	0.014	0.025	0.003
Education quantities	0.208	0.016	0.023	0.003
Other quantities	−0.026	−0.002	0.003	0.000
Total observable prices	−0.049	−0.004	−0.215	−0.024
Education prices	−0.081	−0.006	−0.205	−0.023
Other prices	0.031	0.002	−0.010	−0.001
Unobservables	−0.015	−0.012	−0.007	−0.005
50/10 Ratio				
Total	−0.136	−0.010	−0.050	−0.006
Total observable quantities	−0.018	−0.001	0.007	0.001
Education quantities	−0.017	−0.001	0.039	0.004
Other quantities	−0.001	−0.000	−0.032	−0.004
Total observable prices	−0.119	−0.009	−0.014	−0.002
Education prices	−0.069	−0.005	−0.045	−0.005
Other prices	−0.049	−0.004	0.031	0.003
Unobservables	0.000	0.000	−0.043	−0.002

[a] Changes in the contribution of observable education and observable experience quantities and prices, and unobservables, to the total change in wage inequality.

in both the adjustment and pre-adjustment period, the magnitude of the equalizing education price effect was much larger (on an annual basis) in the pre-adjustment period than in the adjustment period. This evidence is consistent with the interpretation presented in the previous paragraph, that the measured education price effect likely reflects the equalizing effects from the rising supply of education (a trend continuing from the pre-adjustment period) and a small but disequalizing effects of rising relative demand (which shows up only with structural adjustment).

(c) *The source of the difference in inequality outcomes between Chile and Costa Rica*

Summarizing the results of the decompositions: the increase in inequality coincident with structural adjustment in Chile was driven by large disequalizing education price and unobservables effects, while the key to small changes in wage inequality in Costa Rica was an equalizing education price effect. Thus, a key difference contributing to the differences in the distributional outcomes in Chile *versus* Costa Rica were different education price effects. The different education price effects in Costa Rica and Chile were not due to differences in the pace of educational expansion: indeed educational expansion in Chile proceeded at a much more rapid rate than in Costa Rica (Table 3). If supply changes did not cause the different educational price effects, the difference must have been due to changes in the demand for more-educated workers; that is, relative demand shifts for more-educated workers must have been much larger in Chile than in Costa Rica.

Another important conclusion derived from the decompositions is that different inequality outcomes in Costa Rica and Chile were not due to different composition effects of educational expansion. On the contrary, education quantity effects were strongly equalizing in Chile and disequalizing in Costa Rica.

Unobservables effects, which were disequalizing in both countries but much larger for Chile, also contributed to the different inequality outcomes. The pattern of unobservables effects is consistent with the interpretation that the net human capital price effect in Chile was driven by broad skill-biased demand shifts. Juhn *et al.* (1993) argue for the United States that the correlation of the unobservables effects with the observable price effects supports the argument that both effects correspond to the same broad-based skill-biased demand shifts. This argument may be applied to Chile, where unobservables, education prices and other prices were all of the same signs–disequalizing. The argument does not apply, however, to Costa Rica, where unobservables and the education price effects had the opposite signs.

We also estimated the decompositions using the specification of the wage equation used in the original Juhn *et al.* (1993) paper—including complete interactions between the education, experience and region variables. Using this specification, we can calculate total price and quantity effects, but cannot calculate price and quantity effects for individual dimensions of human capital. The results of these estimations are consistent with the results presented in this paper. The signs of the total price, quantity and unobservables effects are the same as those presented above. Using the original technique the measured contribution of unobservables is smaller, and the measured contribution of prices and quantities is larger, than in the decompositions we present here. We believe this suggests that the true unobservables effects are smaller than those estimated in our more disaggregated technique.

(d) *Possible explanations for differences in relative demand changes*

We have shown that an important reason inequality increased in Chile and fell in Costa Rica during their respective periods of structural adjustment was an increase in the relative wages for more-educated workers in Chile compared to relatively stable relative wages for more-educated workers in Costa Rica. Further, we have shown that this increase in relative wages in Chile was driven by an increase in the relative demand for more-educated workers. While the causes of the larger demand shifts in Chile are not the focus of this study, we examine a number of possible interpretations of the different demand shifts in Chile and Costa Rica, including differences between Chile and Costa Rica in the role of: the quality of education, union power, minimum wages, unemployment and macroeconomic cycles, unions, and differences in the speed and comprehensiveness of structural adjustment. [19] In this subsection we examine evidence based on the correlation between and the change in the wage premium for more-educated workers and the possible causes of this change. This evidence is not conclusive, but does strongly suggest a possible explanation for the different changes in relative demand. Specifically, the evidence suggests that the measured returns to schooling for both Chile and Costa Rica do not reflect changes in educational quality, and that union power, minimum wages, and unemployment rates cannot explain the changes in measured returns to education and implied relative demand shifts. We find that the evidence is consistent with the interpretation that trade liberalization in Chile contributed toward increasing returns to education there.

To facilitate the calculations of the correlations, we constructed a single number to measure the wage premium associated with more education by regressing the experience quartic and a single years-of-education variable on the log of real wages. We then used the coefficient on the years-of-education variable as our measure of returns to schooling.

(i) *Quality of education*

Over the period we study, it is possible that there were significant changes in the quality of schooling received by students in Chile and Costa Rica. In Chile, there was a far-reaching educational reform that resulted in, among other things, an increase in the proportion of primary and secondary students attending private schools; in Costa Rica, the 1980s saw an explosion in enrollment in private universities (see Gauri, 1998; Gindling & Berry, 1992). It is conceivable that changes in the quality of education by schooling level over time affect our measures of returns to schooling. If the average quality of schooling by educational level increased over time, the omitted quality variable could have biased our estimates of returns to schooling. To examine this possibility we examined the returns to schooling within narrow age cohorts over time and compared them to our estimates for the entire population of employed male workers. Once workers finish their education, the quality of this already completed education is constant through time. Hence, the distribution of educational quality within an age cohort of workers is nearly constant over time, and the wage structure of a given cohort through time will not reflect changes in the distribution of educational quality. If the returns to schooling within narrow age cohorts are highly correlated with our estimates for the entire population of employed male workers, we can be confident that changes in educational quality are not responsible for the changes in the overall returns to schooling (this is what is done in Katz & Murphy, 1992).

For Costa Rica, the correlation between returns to schooling by age cohorts and the overall returns to schooling for the 1987–95 are positive and large: 0.72, 0.99, 0.81 for cohorts entering the labor force in 1967–72, 1972–77, and 1977–82 for three-year centered averages. For annual point estimates the corresponding correlations are 0.70, 0.95, and 0.6. Turning to Chile we find that estimated returns to schooling within cohorts are also strongly correlated with the overall estimated returns to schooling. In Chile the correlations of estimated returns to schooling by cohorts and overall returns to schooling over 1974–90 are 0.85, 0.88, and 0.84 for three-year centered averages, and 0.85, 0.84, and 0.74 for annual point estimates for cohorts entering the labor force between 1954–59, 1959–64 and 1964–69. We conclude that changes in average educational quality by schooling level do not account for the overall pattern of returns to schooling in either Costa Rica or Chile.

(ii) *Union power*

Because Chile underwent major political changes with the military *coup d'etat* in 1973, it is possible that changes in the relative bargaining power of less-skilled workers influenced relative wages and returns to schooling after 1973.

Specifically, was the rise in relative wages and returns to schooling in Chile over 1974–90 due to a gradual decline in the negotiating power of workers with less than university education (hereafter referred to as union power)? If unions disproportionately represent less-educated workers, or have a policy of negotiating for a reduced wage gap between workers at different education levels, then falling union power would be expected to lower the wages of less-educated workers relative to the wages of workers with university education.

If union power fell gradually over the 1974–90, and especially over 1974–80 when returns to schooling rose most quickly, this gradual decline in union power could potentially explain the measured rise in the returns to schooling. But the pattern of change in union power in Chile over time followed a very different pattern. Instead of a gradual drop in union power after the 1973 *coup d'etat*, union power was almost instantly reduced to zero immediately after the coup, and then *rose* very gradually over 1974–90. Unions and labor laws were strong prior to 1973.

Labor law legislation dated largely from the 1930s and union power reached a peak under the Socialist government of Salvador Allende (1970–73). On 11 September 1973 the pro-union government of Salvador Allende was overthrown in a bloody *coup d'etat* and a right-wing, highly repressive military government under the leadership of Augusto Pinochet, took power. Military control of the country was established overnight. The military's control over the union movement was nearly immediate and absolute. Unions were immediately declared illegal and remained so through 1979. Immediately following the *coup d'etat* on 11 September 1973, union leaders and rank and file union members were rounded up and often summarily executed, imprisoned or exiled, as were political leaders aligned with the union movement. Pro-union political parties were immediately declared illegal and remained so for more than a decade. Severe repression continued through the early 1980s, after which repression became much more limited. Having been declared illegal in 1973, unions regained legal standing in 1979, in the *Plan Laboral*. This legislation legalized unionization and collective bargaining at the plant, though not industry, level. While the legislation was weaker than the pre-1973 legislation, it reflected a gradual increase in union power and increase in union rights compared to 1973–78 when unions were illegal.

In summary, instead of gradually falling over 1974–80, union power plummeted suddenly in September 1973 with the military coup. While never recouping its earlier levels, union power rose over 1973–80 and gradually thereafter until 1990. This gradual rise in union power after 1974 would have tended to exert downward pressure returns to schooling in the period where measured returns to schooling rose. We conclude that the rise in returns to schooling over 1974–90 in Chile was not due to falling union power. [20]

In Costa Rica it is unlikely that wage structure over 1987–95 was much influenced by changing union power. Labor laws were unchanged over this period, and union densities were nearly constant; varying from a minimum of 14% to a maximum of 15%, with no obvious trend. [21] We conclude that wage structure in Costa Rica over this period was little affected by changes in union power.

(iii) *Minimum wages*

The impact of minimum wages upon wages and employment are controversial (see, for example, Card & Krueger, 1995). To the extent that the wages of less-skilled workers are affected by minimum wages while the wages of high-skilled workers are not, a rise in the minimum wage would be expected to increase the wages of less-skilled workers and lower returns to education. To examine the possibility that changes in minimum wages caused changes in returns to schooling, we calculated the correlations between returns to schooling and an index of real minimum wages in both Chile and Costa Rica. In both cases the correlations were insignificant at the 10% level (the correlation coefficients were −0.22 for Chile and −0.43 for Costa Rica—we tested the significance of the correlations by regressing the index of minimum wages on returns to schooling and then testing the significance of the estimated coefficient). [22]

(iv) *Unemployment*

Changes in the level of unemployment might also affect returns to education. If one skill group is disproportionately represented among the unemployed, then their wages would likely fall more rapidly than the wages of the other skill groups, and this could affect the returns to education. Because the skill composition of unemployment can vary, the sign of the potential impact is not obvious *ex ante*. To examine whether the unemployment rate affected returns to education we calculated the correlation between returns to education and unemployment rates in both countries. As with minimum wages, the estimate correlations were insignificant at the 10% level (the correlation coefficients were 0.13 for Chile and 0.36 for Costa Rica—we tested the significance of these correlations by regressing unemployment rates on returns to schooling). [23]

(v) *Structural adjustment and trade liberalization*

Both Costa Rica and Chile implemented structural adjustment programs during the period that we studied, although the structural adjustment program was more rapid and comprehensive in Chile. Arguably the most clear-cut and influential structural reform that Chile undertook after 1973 was trade liberalization. From a long history of import-substitution industrialization, Chile underwent far reaching and rapid trade liberalization after 1973. Average tariffs declined from over 100% in 1973 to 10% in 1980. Nontariff barriers were also eliminated. Tariffs were modestly and briefly increased in 1983 to a peak of 28%, but then quickly lowered again to 11% by 1992. Chile was transformed to an export led economy, with rapid GDP growth, except for the 1982–84 recession provoked by the international debt crisis and recession. The ratio of exports to GDP rose from 11% in 1973 to an average of nearly 30% after 1980. [24]

One line of trade theory based upon the factor endowments approach and the Stolper–Samuelson theorem predicts that trade liberalization raises returns to skill in countries strongly endowed with skill while lowering them in skill poor LDCs. In the Hecksher–Ohlin–Samuelson framework with two labor types, relative wages in the LDC—the ratio of wages to skilled workers over wages of less-skilled workers, monotonically related to the returns to education—are proportional to the relative prices of tradeable goods. In addition, domestic relative prices, the ratio of prices of skill-intensive goods over the prices of unskilled intensive goods, are a mark up over the exogenous international relative price ratio. By the Stolper–Samuelson theorem, lowering tariffs lowers the ratio of tradeable goods prices local producers face, hence lower relative wages and returns to education.

Many trade economists, however, contest the appropriateness of the HOS framework. For example, Bhagwati and Kosters (1994) argue that as economists we need to remind ourselves that the "original" view of the factor price equalization theory was correct: its assumptions are indeed extraordinarily demanding. It is not therefore a compelling, or adequate, guide to real-world phenomena. An alternative model of the impact of trade liberalization upon relative wages, sometimes called Skill-Enhancing-Trade, has been presented (Stokey, 1996). If greater trade accelerates technology diffusion from North to South, then trade can be skill-enhancing. Recent evidence finds that technology transfer occurs via trade (Coe, Helpman, & Hoffmaister, 1997). Trade liberalization may raise relative wages in some LDCs by inducing rapid adaptation of modern skill-intensive technologies from the North. Trade liberalization, and the exchange-rate devaluations that frequently accompany it, increase trade flows. It is well known that real devaluations typically raise the current account surplus, permitting higher levels of machinery imports. Heightened competition from trade liberalization leads to pressures to modernize via importing state-of-the-art machinery. Technology is bundled with this machinery. In LDCs emerging from Import-substitution industrialization strategies that stifled adoption of foreign technologies, this will lead to an initial large jump to more modern and skill-intensive technologies. Subsequently, the liberalized LDC will continue on a skill-intensive biased trend similar to that being observed in the North. Relative wages would follow a similar path conditioned by supply changes in supply. A related hypothesis is that trade liberalization frees up capital flows that will move from the low interest rate, capital-rich North to the high interest rate, capital-poor South. Even without bundled technology, this would lead to higher capital-output ratios. Because of complimentarily between capital and skill, this would raise relative demand for skill (Stokey, 1996).

The evidence for Chile is consistent with the interpretation that trade liberalization affected returns to schooling there, but instead of bearing out the Stolper–Samuelson theorem the evidence is consistent with the opposite view, that trade liberalization led to a rise in returns to schooling. Most convincing, in Chile after 1973 falling tariffs are associated with rising returns to schooling. Examining the correlation of average tariffs and returns to schooling, instead of the positive correlation implied by Stolper–Samuelson, the correlation coefficient between returns to schooling and average tariff levels was negative and large (−0.77). This correlation coefficient is more than twice as large as the correlation coefficients between returns to schooling and any of the other possible explanations examined.

Other evidence is also consistent with the alternative skill-enhancing trade hypothesis in Chile. According to that hypothesis we would expect machinery imports, and associated embodied technical change, to accelerate in the post-trade reform era. Imports of machinery to GDP rose rapidly with trade liberalization, doubling from 3% in 1973–74 to an average of 7% in 1977–80 and then rising further to an average of 10% over 1988–92 (Robbins, 1994). This evidence is consistent with the interpretation that trade liberalization led to an acceleration in imports of physical capital and that the combined capital deepening and likely skill-intensive technology embodied contributed to an increase in the relative demand for skill in Chile. Robbins (1994) shows that the interindustry employment shifts towards more unskilled-intensive sectors that is expected with Stolper–Samuelson did not occur. This work also finds there was sharp within-industry skill upgrading, suggestive of skill-intensive technical change.

In summary, we find that the evidence for Chile is consistent with the interpretation that trade liberalization led to the documented rise in relative demand, and hence returns to schooling there.

Structural adjustment in Costa Rica was more gradual and less comprehensive than in Chile. For example, trade liberalization was implemented by Costa Rican policy-makers in an explicitly gradual manner (Lizano, 1990). Vargas-Alfaro (1993, p. 18) characterizes the 1987–92 period of reform as one of "gradual reduction in nominal and effective protection." The tariff reductions begun in 1987, for example, were allowed to adjust over a period of up to five years. [25] Above we presented evidence that it is likely that trade liberalization was an important cause of the increase in the relative demand for more-educated workers in Chile. Given that evidence, and the slower and less comprehensive nature of trade liberalization in Costa Rica, a logical and reasonable interpretation is that the more rapid and comprehensive trade liberalization in Chile compared to Costa Rica was an important cause of the different patterns of change in the relative demand for more-educated workers.

Thus we have found that the returns to schooling in Chile were strongly, negatively, correlated with average tariff rates. A possible causal explanation linking falling tariffs and rising relative demand and returns to education is that trade liberalization accelerated imports of physical capital and skill-biased technological change. Consistent with this view, machinery imports grew rapidly with trade liberalization in Chile. Examination of evidence for key competing explanations, changing educational quality, union power, minimum wages and the unemployment rate appear unable to explain the pattern of returns to education and relative demand. Thus, we find that a logical and reasonable interpretation of the different returns to schooling and relative demand shifts for Chile and Costa Rica lies in the much more rapid and far-reaching trade liberalization in Chile.

5. CONCLUSION

Despite many similarities, during their respective periods of structural adjustment inequality increased much more rapidly in Chile than in Costa Rica. This paper examines the causes of this difference by decomposing changes in inequality into price, quantity and unobservables effects, where price and quantity effects are further separated into education and experience effects. Our first finding is that differences in education price effects were an important contribution to the different inequality outcomes. Changes in education prices were strongly disequalizing in Chile while largely neutral in Costa Rica. Our second finding is that education quantity effects associated with the pace and pattern of education expansion were not the cause of the different inequality outcomes. On the contrary, education quantity effects were equalizing in Chile and disequalizing in Costa Rica. Our third finding was that disequalizing effects of unobserved variables in Chile contributed substantially to the different inequality outcomes.

Further, we found that the different education price effects were not due to differences in the rate of educational expansion in Chile and Costa Rica; educational expansion was more rapid in Chile than in Costa Rica. If supply growth alone had driven the education price effects, we would have seen equalizing education price effects in both Chile and Costa Rica. Therefore, the different education price effects must have been due to more rapid changes in the relative demand for more-educated workers in Chile compared to Costa Rica. The more rapidly rising returns to education in Chile compared to Costa Rica, combined with the more rapid increase in educational expansion in Chile, imply that relative demand for higher education rose more rapidly in Chile than in Costa Rica.

A full exploration of the causes of the more rapid relative demand shift in Chile exceeds the scope of this study. We provide evidence, however, supporting the interpretation that the more rapid rise in relative demand for more-educated workers in Chile during the period of structural adjustment was partly due to the more rapid and radical trade liberalization undertaken there; returns to schooling in Chile were strongly, negatively, correlated with average tariff rates. One possible reason why falling tariffs could lead to rising relative demand is that trade liberalization accelerated imports of physical capital and skill-biased technological change. Consistent with this view, machinery imports grew rapidly with trade liberalization in Chile.

NOTES

1. In 1990, GNP per capita was US$1,940 in Chile and US$1,900 in Costa Rica (World Bank, 1992).

2. "Relative wages" may be understood as the ratio of the wages of more to less-educated workers. Similarly

"relative supply" and "relative demand" are, respectively, the ratio of the supply and demand of more to less-educated workers. See Katz and Murphy (1992), Robbins (1994), and Robbins and Gindling (1999) for theory, precise definitions, and results.

"Relative wages," "returns to education," "the wage premium for more-educated workers," and the "education-price effects" are terms we will use in this paper for closely related concepts whose measures are typically closely, positively, correlated.

3. For additional references to the effects of educational expansion on the distribution of wages, see Robinson (1976), and Almeida dos Reas and Paes de Barros (1991).

4. The Bourguignon *et al.* (1998) and Fields (1998) techniques are briefly described in Appendix A.

5. It is important to note that there are important discontinuities in some variables during 1985–87. Between these two years the design and sample of the surveys changed, raising questions about the comparability of the pre-1987 data with the post-1986 data. Among the unrealistic changes reported from the 1986 to 1987 surveys is a decline in average education levels. The post-adjustment surveys, 1987–95, are comparable to each other.

6. In Chile, we consider structural adjustment to begin in 1974, one year after the military coup. We begin in 1974 rather than 1973 in part because 1974 is the year the variance of earnings is lowest in Chile. While it would also have been reasonable to begin the analysis of structural adjustment in Chile in 1973 or 1975, the results in this paper do not differ qualitatively if we used those years. We chose to end the period of structural adjustment in Chile in 1990. By 1990 the structural adjustment reforms had had nearly 15 years to take effect. Also in 1990 Chile returned to civilian rule and began of new economic reforms that likely also affected the distribution of earnings.

We begin the structural adjustment program in Costa Rica in 1987 for several reasons. First, while some structural adjustment in Costa Rica began in 1984, we lack survey data for 1984 and 1986. Second, changes in survey design impede reliable comparisons between 1985 and subsequent years; there are important discontinuities in the variables in the Costa Rican household surveys during 1985–87. Finally, structural adjustment, notably trade reform, only began in earnest in this year, and was implemented very gradually (Gindling & Berry, 1992, 1993). While legislation for the first structural adjustment program was passed in 1984, many of the most important reforms did not begin for several years. In particular, tariff reductions did not begin until October 1987 and were implemented over five years.

The results presented in this paper are not qualitatively different from those obtained if we compare Costa Rica in 1983 or 1985 with 1995, or if we compare Chile in 1973, 1975 or 1976 with 1990. Tables similar to Table 1 comparing 1973, 1975 and 1976 with 1990 in Chile, and tables comparing 1983 and 1985 with 1995 in Costa Rica, are available from the authors.

7. To facilitate the Chile–Costa Rica comparison, we use the same scale on the vertical axis for both countries.

8. Depending on what year one uses as the start of structural adjustment in Costa Rica, and on the measure of inequality, inequality either increases or decreases slightly in Costa Rica during structural adjustment. If we compare 1983 or 1985 to 1995, then the variance increases, if we compare 1987 to 1995, then the variance falls. (A previous footnote discusses why it might be reasonable to begin the analysis of structural adjustment in Costa Rica with either 1983, 1985 or 1987.) During 1987–90 the 90/10, 90/50 or 50/10 ratios of hourly wages all increase (Figure 2). For the purposes of the Chile–Costa Rica comparison, whether inequality falls or rises in Costa Rica with structural adjustment is less important that the fact that any increase in inequality in Costa Rica with structural adjustment is much less than in Chile, which is true no matter what measure of inequality is used, or what year is considered as the start of structural adjustment.

9. In this paper we consider only the distribution of wages; we have no data on changes in the distribution of property or capital income in Costa Rica and Chile. Property or capital income is likely to be more unequally distributed than labor income.

10. To calculate the average (over-all-years) distributions and coefficients we use all years of data available to us in each country. These years are 1957–95 for Chile and 1976–95 for Costa Rica.

11. Following Juhn *et al.* (1993), we estimate wage equations with education dummies for primary school, incomplete high school, complete high school and university, and a quartic for experience (plus a dummy variable for region in Costa Rica). Unlike Juhn *et al.* (1993), we do not include interaction terms between experience and education. While excluding the interaction terms makes the functional form less flexible than that used in Juhn *et al.* (1993), using the less flexible functional form permits us to estimate quantity and price effects separately for education and experience.

12. In calculating the decompositions, we first change the quantities and prices associated with education, and second change the quantities and prices associated with experience. An alternative "path" would be to first change the quantities and prices associated with experience, and second change the quantities and prices associated with education. As Bourguignon *et al.* (1998) note, the results of the decomposition may differ depending on the path used.

13. For example, to calculate the contribution of other quantities to the change in the variance we first subtract the variance of Y_{it}^2 from the variance of Y_{it}^1, and then report the changes in this difference over time. To calculate the contribution of other quantities to the change in the 90/10 ratio we subtract the 90/10 ratio of Y_{it}^2 from the 90/10 ratio of Y_{it}^1, and so on.

14. The sum of the contribution of educational quantities and other quantities equals the contribution of "total observable quantities."

15. The sum of the contribution of educational prices and experience prices equals the contribution of "total observable prices."

16. Tables in Appendix B present regression results for the pooled regression, and regression results and the means and S.D. of the variables used in the regressions for the beginning and end years of the analysis (1974 and 1990 in Chile and 1987 and 1995 in Costa Rica). In the pooled regressions, all coefficients are significant. In the regressions for the individual years, the coefficients on the education dummy variables are always significant while only a sub-set of the experience variables are significant in each year.

17. We re-calculated the decompositions using the other possible "path" (see note 12)—first changing the distribution of experience holding the distribution of education constant, and second changing both the distribution of experience and education. The results using this path are similar to those presented in this paper, both in signs and magnitudes.

18. Table 5 presents changes during 1976–85 in Costa Rica. We do not compare 1976 with 1987 because discontinuities in the survey variables between 1986 and 87 make the pre-1987 surveys non-comparable to the post-1986 surveys. The surveys during 1976–85 are comparable to each other, as are the surveys during 1987–1995. We cannot report results using the 1986 survey because the 1986 data tapes do not include data on years of education completed, and therefore cannot be used to calculate the decompositions.

19. We would like to thank three anonymous referees for pointing out these possible interpretations.

20. The subsection on union power draws heavily from Barrear and Valenzuela (1986).

21. Data on union densities from a personal communication from Juan Diego Trejos, who based the estimates on information provided by the Costa Rican Ministry of Labor.

22. Data on minimum wages are from the Costa Rican Ministry of Labor, Division of Minimum Wage and the Chilean National Statistics Institute.

23. Data on unemployment rates are calculated from the household surveys.

24. The data cited above come from the World Bank (1998).

25. Robbins and Gindling (1999) present a more detailed description of the process of trade liberalization in Costa Rica.

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APPENDIX A. THE BOURGUIGNON *ET AL.* (1998) AND FIELDS (1998) TECHNIQUES

(a) Fields (1998) presents an inequality decomposition technique based on the factor decomposition of Shorrocks (1982). He derives an exact decomposition of a general class of inequality measures into the share of total income inequality explained by measured income-determining characteristic. The share of income inequality explained by each income-determining characteristic j, $s_j(\ln Y)$, is

$$s_j(\ln Y) = [B_j * \mathrm{std}(X_j) * \mathrm{cor}(X_j, \ln Y)]/\mathrm{std}(\ln Y), \tag{A.1}$$

where Y is income/earnings, B_j is the coefficient on income determining characteristic X_j.

For infinitesimal changes in X_j, Fields (1998) also develops an expression for further decomposing the change in the contribution of each income determining characteristic into price and quantity effects.

$$s_j(\ln Y) = B_j + \mathrm{std}(X_j) + \mathrm{cor}(X_j, \ln Y) - \mathrm{std}(\ln Y), \tag{A.2}$$

where the italics indicate changes. In Eqn. (A.2), B_j is the price effect associated with income determining factor j, and $\mathrm{std}(X_j)$ is the quantity effect associated with income determining factor j. As Fields notes, the decomposition described in (A.2) is valid only for small changes in each income determining characteristic j. Fields (1998) also notes that "An objection can be raised... that B_j and $\mathrm{cor}(X_j, \ln Y)$ are both functions of $\mathrm{cov}(X_j, \ln Y)$, so that one cannot be varied without the other" (p. 19).

In the Fields (1998) decomposition technique, individual price and quantity effects can only be calculated for single variables. For example, one cannot calculate the impact of changes in education prices on inequality if the education variable is specified as a set of dummy variables (as it is in our paper). One can, however, calculate the education price and quantity effects if one specifies the education variable as a single years-of-education variable.

Gary Fields and Jesse Leary generously provided us with the STATA programs needed to calculate the Fields (1998) decompositions.

We did so, using a quartic in experience and a single years-of-education variable as right-hand-side variables. The results from these decompositions are consistent with the decomposition results presented in the body of our paper. These results are that (i) the increase in inequality in Chile after 1974 was due to increases in the relative wage paid to more-educated workers (the "education price effect"); (ii) the education quantity effect in Chile was equalizing; (iii) in Chile the effect of changes in experience prices and quantities was small; and most importantly (iv) the different inequality outcomes between Chile and Costa Rica-an increase in Chile and a decrease in Costa Rica—were due to the different education price effects. The education price effects were equalizing in Costa Rica and, as noted above, disequalizing in Chile. In Costa Rica, education quantity, experience quantity and experience price effects are all disequalizing.

(b) The technique that we derive in the body of this paper decomposes inequality in the distribution of personal hourly wages. Bourguignon *et al.* (1998) present an inequality decomposition of the more complete measure of household income. The Bourguignon *et al.* (1998) technique is based on the estimation of earnings equations corrected for selectivity bias using the Heckman two-step "Mills ratio" procedure. Bourguignon *et al.* (1998) correct for two types of selectivity: selection into the labor force and selection into salaried employment or self-employment. Like the decomposition that we present in the present paper, the Bourguignon *et al.* (1998) technique uses the estimated earnings equations to simulate the distribution of income allowing the price or quantity of one income-determining characteristic to change sequentially, while holding other price and quantities constant. Since many income-determining characteristics, such as education, influence the decision to participate in the labor force as well as the structure of earnings, the impact of variables such as education on household income inequality works through two mechanisms: the impact on the labor force participation decision (and therefore on the number of household members who are working) and the impact on the structure of earnings for those who do work. Using selectivity-corrected earnings equations allows Bourguignon *et al.* (1998) to decompose the price and quantity effects into the effects on participation and the effects on the structure of earnings.

The use of the Heckman procedure to correct for selectivity bias, upon which the Bourguignon et. al. methodology is based, has been criticized for not being robust to violations in underlying assumptions. As Kennedy (1998) writes

> Monte Carlo studies....find that on a MSE criterion, relative to subsample OLS the Heckman procedure does not perform well when the errors are not distributed normally, the sample size is small, the amount of censoring is small, the correlation between the errors of the regression and the selection equations is small, and the degree of collinearity between the explanatory variables in the regression and selection equations is high. It appears that the Heckman procedure can often do more harm that good, and that subsample OLS is surprisingly efficient, and more robust to non-normality (p. 156).

(*Appendix B—continued overleaf*)

APPENDIX B

Table 6. *Means and S.D. of the explanatory variables used in the earnings equations: Chile (1974–90) and Costa Rica (1987–95)*[a]

	Chile		Costa Rica	
	1974		1987	
	Mean	Standard Deviation	Mean	Standard Deviation
Education dummies				
Primary incomplete	0.229	0.421	0.301	0.459
Primary complete	0.323	0.468	0.356	0.479
Incomplete secondary	0.129	0.335	0.141	0.348
Complete secondary	0.112	0.315	0.123	0.328
University	0.207	0.405	0.078	0.267
Experience	20.85	13.09	17.39	13.39
Central valley dummy	na	na	0.446	0.497
	1990		1995	
Education dummies				
Primary incomplete	0.078	0.268	0.239	0.427
Primary complete	0.156	0.363	0.380	0.485
Incomplete secondary	0.178	0.382	0.156	0.363
Complete secondary	0.277	0.448	0.111	0.314
University	0.311	0.462	0.114	0.318
Experience	18.31	12.62	18.29	13.09
Central valley dummy	na	na	0.434	0.495

[a] Data from male salaried employees only.

Table 7. *Earnings equation results from regressions pooling data sets: Chile and Costa Rica*

	Chile		Costa Rica	
	Coeff.	Standard Error	Coeff	Standard Error
Education dummies				
Primary complete	0.248*	0.000	0.148*	0.005
Incomplete secondary	0.513*	0.009	0.367*	0.001
Complete secondary	0.925*	0.009	0.716*	0.002
University	1.607*	0.009	1.261*	0.007
Experience	0.086*	0.002	0.019*	0.0014
Experience^2	−0.0026*	0.0002	0.0007*	0.00008
Experience^3	0.000035*	4.0E−7	−0.00003*	0.0000001
Experience^4	−1.9E−7*	1.0E−9	−2.1E−7*	1.0E−9
Central valley dummy	na	na	0.034*	0.004
Constant	−6.136*	0.014	0.997*	0.007
R-squared	0.396		0.325	
Observations	70444		89942	

* Significantly different from zero at 1%.

Table 8. *Earnings equation results: Chile (1974–90) and Costa Rica (1987–95)*

	Chile		Costa Rica	
	1974		1987	
	Coeff.	Standard Error	Coeff.	Standard Error
Education dummies				
Primary complete	0.195*	0.090	0.124*	0.021
Incomplete secondary	0.359*	0.050	0.335*	0.027
Complete secondary	0.713*	0.050	0.636*	0.028
University	1.227*	0.046	1.204*	0.033
Experience	0.068*	0.016	0.0027	0.0054
Experience^2	−0.0016	0.001	0.0017*	0.00034
Experience^3	0.000013	0.000029	−0.00005*	0.0000007
Experience^4	−6.0E−8	7.0E−8	3.28E−7*	5.0E−8
Central valley dummy	na	na	0.083*	0.016
Constant	−6.014*	0.084	1.154*	0.025
R-squared	0.347		0.315	
Observations	1822		5074	
	1990		1995	
Education dmmies				
Primary complete	0.142**	0.074	0.061*	0.019
Incomplete secondary	0.260*	0.075	0.256*	0.022
Complete secondary	0.667*	0.073	0.515*	0.026
University	1.550*	0.073	1.093*	0.026
Experience	0.128*	0.018	0.020*	0.005
Experience^2	−0.0064*	0.0013	0.00043	0.00035
Experience^3	0.00014*	0.000035	−0.00002*	0.00000087
Experience^4	−1.1E−6*	3.0E−7	1.6E−7**	6.9E−8
Central valley dummy	na	na	0.095*	0.014
Constant	6.163*	0.101	1.218*	0.023
R-squared	0.403		0.334	
Observations	1811		6459	

* Significantly different from zero at 1%.
** Significantly different from zero at 5%.