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Explaning Earnings and Income Inequality in Chile

Grisha Alexis Palma Aguirre

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Grisha Alexis Palma Aguirre

To Elliot, Gabriella, and to my beloved wife Annika

Abstract

The focal point of all papers in this thesis is income inequality in Chile. In some of them household income is analyzed, in others monthly earnings or the wage rate are used. In the first and fourth paper a long-run analysis is done, while in the second and third I concentrate my attention only on the 1990s. Despite the different periods covered, the variables analyzed, and methods used, in general the thesis attempts to bring to light the factors that help to explain the levels and changes of income inequality in Chile, a country that has a marked position among the most unequal in the region and has gone through great political and economic changes in the last decades. Moreover, Chile posses some surveys of fairly good quality that permit long-run or country-wide analysis of the distribution of earnings and income. Therefore, Chile is not only an interesting country to study, but it is also encouraging.

The first paper summarizes the development of household income and earnings inequality of several groups of the Chilean labor market during the period 1958-2004. Furthermore, it surveys the explanations given in existing studies to the levels of and trends in income inequality in the last decades. With the exception of the early 1970s, income inequality deteriorated from 1957-63 to the 1987-90 period, before declining in 1991-98. The rate of return to university education, as well as the dispersion of hourly earnings of males and white-collar workers, has generally followed the overall pattern of inequality. Education was found to be a key factor behind the dispersion of incomes and earnings in Chile, explaining 13-40% of the inequality, depending on the survey, definition of income, year, method, and sample used. Openness and trade has been suggested to be important to understand the deterioration of Chile's distribution of wages through the increased demand for skilled workers that followed the external sector liberalization after the mid 1980s. Evidence that increased female labor-force participation and higher rates of unemployment have an inequality increasing effect is provided by various studies.

The second paper focuses on several important but relatively unexplored issues in the body of relevant literature. Using a Bootstrap technique, and analyzing self-employed workers separately from employees, this paper presents several interesting results. Wage and salary and household income inequality deteriorated significantly at the end of the decade while the dispersion of self-employment incomes deteriorated

in the 2000-03 period with respect to 1992 and 1996-98. Accounting for 20%-40%, the between-group component of education was an important factor to explain the level of inequality. But as income disparities between educational groups grew larger during the 1990s, education played a dominant role to explain even the change of wage and salary and self-employment income inequality. This was corroborated analyzing the Gini coefficient of household income by source, which reveals that its underlying components changed in size although the Gini remained at 0.54 all the three years analyzed. The earnings of employees and the self-employed with university education accounted for over 26% of the Gini coefficient of household income in 1990. By 2003, this share had gone up to 40%. The contribution of earnings of primary educated employees and self-employed workers declined from 7% to less than 2% over the same period.

The third paper studies the distributional effects of an occupational change that occurred in Chile's labor market from 1992 to 2000. During this period the employment structure shifted towards informal employment (from 9% in 1992 to 15% in 2000), but also towards professional occupations (19% in 1992 to 26% in 2000). Both within-group and between-group composition-effects increased inequality, while within-group and between-group change in variance reduced it. However, using the inequality-decomposition of Fields and Yoo (2000) to analyze the inequality within-occupations it is found that even education and hours-worked had an increasing effect in overall inequality, while all the other variables in the earnings equation, and especially the residuals, had an inequality-decreasing effect.

The fourth paper analyzes in detail the inequality of hourly earnings of male workers in Santiago. Analyzing the years 1974, 1987, 1992, 2000, and 2003 I concentrate my attention on the extreme values in inequality of the last decades. Using an Oaxaca-Blinder type decomposition, but implementing quintile regressions, I am able to decompose changes in the male wage inequality into a price effect, a composition effect, and a residual. The first important finding is that inequality in the upper part of the distribution seems to be more important than the dispersion in the lower part to total inequality. Second, the large deterioration of male wage inequality between 1974 and 1987 was the result of the combined price effect and composition effect. For other inequality changes the price effect was dominant. Among the variables included in the quintile wage equations, such as age, education, occupation, and sector,

education was the one with the largest composition effect.

Keywords: Income Inequality, Chile, Bootstrap, Quintile Regressions.

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that if this has been a dificult time for me, it has been even more difficult for you.

Gothenburg, December 2007
Alexis Palma

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Introduction and Summary

1 Introduction

This thesis is made of four different essays. Each one of these papers, but using different methods, attempt to contribute to an increased understanding of the factors that determine the level and pattern of Chile's earnings and income inequality. The reasons for my interest in this country are the following: First, Chile has drawn increased attention during the last 15 years for its ability to deliver high rates of GDP-growth in a context of macroeconomic stability. Along with a rapid economic expansion, increased exports, and declining unemployment rates, poverty rates halved between 1990 and 2000. On the other hand, the distressingly high level of income inequality left by the military junta, declined only during some few years and even deteriorated during the second half of the 1990s. In consequence, Chile maintained its relative high position among the most unequal countries in the region. Therefore, income inequality issues started to draw increased attention of scholars who questioned themselves why Chile's unequal dispersion of income seemed so difficult to decline. The extremely uneven pattern in Chile's distribution of income became regarded as one of the most serious problem facing Chilean policymakers (Hojman, 1996). Therefore, when the seminal paper of Atkinson: *Bringing Income Distribution in from the Cold* (Atkinson, 1997) was published, income inequality was already a hotly debated issue in Chile. This thesis attempts to make a small contribution to this on-going discussion in Chile.

Second, Chile has several data-sets of fairly good quality, two of which are used in this thesis, that allow long-run or country-wide analysis of earnings and income inequality. However, these data have some weakness that we should be aware of. The Employment and Unemployment Survey for Santiago is only collected in the Metropolitan region, therefore excluding information about provincial and rural households. On the other hand, this survey covers four decades during which Chile went through great economic and political changes. *Caracterización Socioeconómica Nacional* (CASEN) has the advantage of being representative for the whole country, but it is only available for 1987, 1990, 1992, 1994, 1996, 1998, 2000, and 2003, of which the 1990-2003 versions are used in this thesis.

2 Summary of Essays

The first paper presents own calculations of the Gini coefficient of earnings of several segments of the labor markets and provides a summary of the relevant literature. Other methods applied in my thesis are a decomposition of the Theil index (of wages and salaries and self-employment incomes) by population sub-group and a decomposition of the Gini coefficient by income source of household incomes. In the third paper I use a decomposition of observed inequality changes in a between and a within-group component to analyze the effect of a change in the structure of the labor market between 1992 and 2000. The method used in the fourth paper is an Oaxaca-Blinder type decomposition with quintile regressions. I analyze the effect of changes in the composition and in the price of labor market characteristics on the inequality changes observed between 1974 and 1987, between 1987 and 1992, between 1992 and 2000, and between 2000 and 2003.

The main result of paper one is that household income, white-collar, as well as male earnings inequality increased significantly between the early 1970s and the late 1980s. In adition, the inequality between white- and blue-collar workers, and between employers and blue-collar workers deteriorated during the same period. This indicates that not only inequality within but also between labor market groups increased significantly during the 1980s. The survey of the available literature on inequality in Chile indicates that education is the single most important factor behind the distribution of earnings and income, accounting for between 13% and 40% of inequality, depending on the survey, definition of income, year, method, and sample used. The second most important variable is occupation, accounting for 8%-24%, which is even more important than education in the analysis of Amuedo-Dorantes (2005). Intertemporal analysis suggests that a more liberalized external sector, higher rates of unemployment, and higher rates of women participation in the labor market tend to increase inequality in the case of Chile.

The results of the second paper, using national representative data, reveal that wage and salary and household income inequalities were significantly higher at the end of the decade than in 1994. The distribution of self-employment income was more unequal in the period 2000-03 compared to 1996-98. Although wages and salaries and self-employment incomes followed a somewhat different pattern, underlying income disparities between educational groups grew larger during the 1990s for both these income variables. In consequence, the between-education component of inequality changes was positive, large, and significant. The analysis of the Gini coefficient of household income of 1990, 1996, and 2003 provides evidence of a changing structure of inequality with a 50% increase in the share of inequality explained by the earnings of employees and self-employed with university education. The contribution of labor income of employees and self-employed with primary education became less and less important during the research period. In essence, this analysis indicates that behind the relatively stable pattern of inequality during the 1990s, the underlying structure of inequality changed over time.

The focal point of the third paper are the years 1992 and 2000. Between these years I found a change in the structure of the labor market with a growing share of informal employment (from 9% to 15%) and professional occupations (from 19% to 25%). Applying a technique developed by Fields and Yoo (2000) I decompose the inequality change between these years in a within-group and a between-group component, which in turn are decomposed into a composition- and a change in variance-effect. The relatively small inequality increase observed between 1992 and 2000 was the result of an inequality-decreasing effect of the change in variance within-groups and an inequality-increasing effect of the other components. In particular, the informal sector had a large and inequality-increasing contribution in the inequality decomposition.

The analysis of male wage inequality in Santiago in the years 1974, 1987, 1992, 2000, and 2003 is based on an inequality decomposition developed by Machado and Mata (2004). The pattern of the Log-wage inequality, and the inequality in the upper part of the distribution, are characterized by a significant growth between 1974 and 1987 and a great decline between 1987 and 1992. In the following years inequality was more stable than in previous decades but was characterized by an inverted-U pattern between 1992 and 2003. I found significant changes in inequality through consecutive comparisons of 1974, 1987, 1992, 2000, and 2003, which represent extreme values in the distribution of wages during the analyzed period. An underlying reason of the great deterioration in inequality between 1974 and 1987 was a significant price and a composition change, both of which had an inequality increasing effect. The following inequality decline, between 1987 and 1992, is mainly explained by the significant inequality decreasing effect of prices, which explains the totality of the first and fifth deciles' change between these years.

References

Amuedo-Dorantes, C. (2005): "Work Contracts and Earnings Inequality: The Case of Chile," *Journal of Development Studies*, 41:4, pp. 589-616.

Atkinson, A. B. (1997): "Bringing Income Distribution in from the Cold," *The Economic Journal*, 107, pp. 297-321.

Hojman, D. E. (1996): "Poverty and Inequality in Chile: Are Democratic Politics and Neoliberal Economics Good for You," *Journal of Interamerican Studies and World Affairs*, 38(2/3), pp. 73-96.

Fields, G., and G. Yoo (2000): "Falling Labor Income Inequality in Korea's Economic Growth: Patterns and Underlying Causes," *Review of Income and Wealth*, 46:2, pp. 139-160.

Machado, J. A. F., and J. Mata (2004): "Counterfactual Decomposition of Changes in Wage Distributions Using Quintile Regression," *Journal of Applied Econometrics*, 19, pp. 1-21.

Income Inequality in Chile: Levels, Trends, and Explanations

ALEXIS PALMA [†‡]

Abstract

The primary concern of this paper is the inequality of household income and earnings across several segments of the Chilean labor market from 1958-2004. Accordingly, I examine the levels of and trends in income inequality and analyze the explanations for their existence. With the exception of the early 1970s, income inequality deteriorated from 1957-63 to the 1987-90 period, before declining in 1991-98. The rate of return to university education, as well as the dispersion of hourly earnings of males and white-collar workers, has generally followed the overall pattern of inequality.

Education is a driving factor of the concentration of income in Chile. Depending on the survey, definition of income, year, method, and sample used, education accounts for up to 40% of Chilean income inequality. Additionally, because it generated a demand change favorable to skilled workers both between and within sectors, openness to international trade has played an important role in increasing inequality.

Higher unemployment rates and higher participation of females had an inequality increasing effect, while higher minimum wages reduced the degree of inequality. Public policies have also been important to improve the distributive situation during the 1990s, when they increased the income share of the first quintile considerably.

[†]Department of Economics, Göteborg University, Box 640, SE-405 30 Göteborg, Sweden. Phone: +46 (0)31 786 13 67. E-mail: alexis.palma@economics.gu.se

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1 Introduction

Issues pertaining to income inequality have attracted increased attention in recent years; both theoretical and empirical studies have examined cross-country differences and within-country changes in the dispersion of income. One clear result emerging from these studies is a severe critique of the Kuznets' hypothesis that there exists an inverted-U relationship between the level of economic development and income inequality. Empirical studies have found only weak support for this hypothesis (Fields, 2001), and some data have even indicated an opposite, U-shaped relationship. These conflicting findings suggest that new insights are needed to understand the main factors behind the level of inequality and its change over time. Because income distribution is the result of numerous decisions made by households, firms, organizations, and the public sector, and of micro- and macro-level economic shocks, both in-depth and broad analyzes of individual countries are required. This paper is therefore an attempt, utilizing two primary methods, to summarize Chile's income inequality pattern over the last five decades: original calculations and analysis of Gini coefficients and, in order to explore the explanatory factors of income inequality, a survey of recent studies on the topic.

The underlying reasons for my interest on Chile are mainly three. First, Chile posses a survey that, although only concentrated to the Metropolitan region, it covers more than four decades and have experienced relatively few changes over the years with respect to questions included and methodology. Second, Chile has undergone significant economic, political, and income inequality changes over the last 40 years. During the early 1970s, under a left-wing regime, Chile's economy was subject to increased intervention of the government, a widespread socialization, and increased redistributive policies. In consequence, inequality figures from Santiago fell to historically low levels. By the 1980s, under the control of a military junta, the economy become much more outwardly oriented, liberalized, but also portrayed by great fluctuations in output and in the dispersion of incomes. By the end of the 1980s, the Gini coefficient of household income had reached a level that was 20% higher than in the early 1970s

During the 1990s, when the country returned to democracy, the Chilean economy was characterized by sustained growth and considerable socioeconomic progress as real GDP per capita increased by more than 100% from 1985 to 2000. In this context of rapid economic expansion, the national incidence of poverty declined from 40% in 1987, to just 17% in 2000. Also in terms of income distribution improvements were achieved but only during some few years. After 1994 inequality seems to have stagnated and even deteriorated at the end of the decade.

Third, despite that the Chilean economy delivered high figures of GDP-growth for most of the 1990s and the country has climbed to one of the richest of the region, Chile has maintained its position as one of the most unequal countries in Latin America. In fact, a study by the World Bank (World Bank, 2003) shows that, using national representative data, Chile was the fourth most unequal country in Latin America in the early 1990s; by the end of the decade, only Brazil had a higher concentration of household income. As a result, many policymakers and scholars, along with the wider Chilean public, have pointed to the failure to reduce inequality as a major drawback in the country's otherwise successful economic development of the last 15 years. In consequence, scholars have generated an increasing amount of research to explain the high level of inequality in Chile. Therefore, this paper is an attempt to summarize Chile's pattern of inequality during the last decade and the results of the studies that try to explain it.

The first part of this paper examines, in several periods between 1957 and 2004, the levels of and trends in inequality of household income and earnings in several segments of the Great Santiago labor market. The relevant dataset is one of only a few that have largely maintained their original format over several decades. Prior to this study, Larrañaga (2001) performed a similar analysis of inequality across different political regimes in Chile; however, I examine the income inequality of specific groups not included in his study, such as white-collar workers, blue-collar workers, males, females, and others. In this way I provide a broader picture of inequality than the one provide by Larrañaga. Moreover, I study the significance of inequality changes using a Bootstrap technique developed by Mills and Zandvakilly (1997). The second part of this paper reviews several explanations of income inequality in Chile that concentrate on the four areas most relevant to the issue: the role of education, the external sector liberalization, the labor market, and public policies.

The next section presents detailed information on the size of and change in the Gini coefficient for different groups and periods between 1957 and 2004 in the Great Santiago area. Section three explores the distributional effects of education, openness to trade, the labor market, and public policy. Finally, in the fifth section, I summarize the issues discussed in this paper and present conclusions.

2 Income and Earnings Inequality Trends

2.1 Levels and Trends in Great Santiago

Using data from the Employment and Unemployment Survey of Universidad de Chile for Greater Santiago (EUSS), this section presents an original calculation and analysis of the Gini coefficient of the household incomes per capita and hourly earnings of several segments of the labor market.

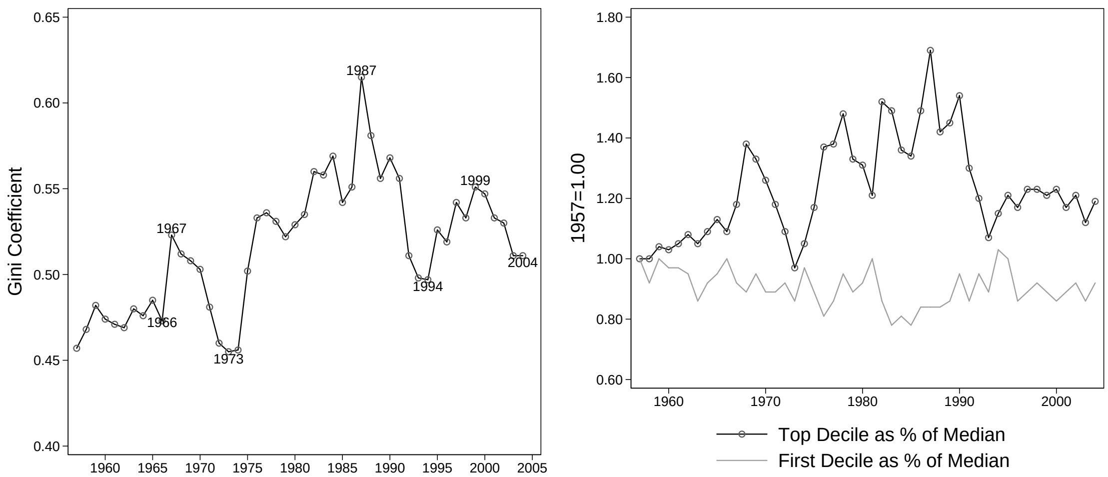
EUSS is an annual cross-sectional household survey of 10,000 individuals in Great Santiago. The strength of the survey lies in its relatively unchanged format with respect to questions and sampling methodology, an excellent characteristic for income inequality analysis over long periods of time. Additionally, in order to maintain the representative nature of the survey, its sampling process has been reviewed several times over the years (Contreras, 2002). An important weakness of EUSS is its geographical concentration; as it is carried out only in the Great Santiago area, the survey contains no information pertaining to rural and provincial households.[1]

During the 1960s, Chile introduced land and job-security reforms and took the first steps toward nationalizing its copper industry in an environment of political, social, and labor-based activism. In the face of 25% average inflation from 1964-70, real GDP grew at 4% per year, including an extraordinary high growth rate, 11%, in 1966. In this context of moderated and relatively stable GDP-growth, unemployment declined reaching 5% in 1970. Moreover, as policy makers introduced a new policy of 100% wage indexation, on past inflation, real salaries increased by 8% per year. The unusual high rate of GDP-growth is probably the reason of the marked increase of the Gini coefficient in 1967 when this measure of inequality increased by 10% after several years of stable inequality. However, after a closer inspection of the distribution of household income it becomes evident that inequality behaved very differently in the lower and upper part of the distribution. The top decile, as percentage of the median, increased marginally until 1967 when it jumped to 1.40. Since the bottom decile did not reported any clear trend, the deterioration of inequality in 1967 was mainly driven by the inequality in the upper part of the distribution.

[1] For more detailed information regarding these data, see Palma (2006).

Figure 1

Gini Coefficient of Household Incomes per Capita and Inequality at the Different Parts of the Distribution in Great Santiago, 1957-2004



Source: Author's calculation from EUSS, 1957-2004.

Under the left-wing regime that came to power in the 1970s, the state gradually became a more important economic player; the government not only increased aggregate demand, but, by nationalizing several industries, it also gained control over important sources of output (Larrain and Meller, 1991). The first two years of the Allende administration were successful in several aspects. The pre-Allende GDP-growth rates continued in 1970 and even escalated to 8% in 1971. Also inflation continued at pre-Allende rates while unemployment declined to historically low levels, 3%, by the 1971-72 period. In real terms, average wage increased by 23% in 1971, with the minimum wage of blue-collar workers increasing at 39%, far above that of other labor market groups, 10%. In 1972 the situation deteriorated considerable with the inflation rate climbing to over 200% and GDP-growth and the growth rate of wages turning negative. The strong orientation towards redistributive policies of the government in place is reflected in Figure 1 by a clear reduction of the degree of inequality in 1971 and 1972, in both cases by 20 Gini points. During this period the bottom decile continued to fluctuate without dramatic changes. The patterns is clearly different for the top decile, which declined by nearly 30% between 1968 and 1972.

In the 17 subsequent years of military regime, Chile turned 180 degrees toward an orthodox market economy, generating tremendous change in several sectors and institutions of the economy. Price controls were eliminated, but inflation reduction became a top priority for the new administration. Tariffs were unified and lowered, union activity was banned, and the public sector was dramatically reduced; some of the land expropriated during the land reform was either returned to its former owners or sold. In addition to dramatic political changes, great economic fluctuations characterize this period in Chile's history. The first great recession took place in 1975, when GDP declined by more than 12%. In 1982, a combination of domestic- and external-sector imbalances caused a second devastating economic crisis. As output fell by 14%, the unemployment rate climbed to more than 30% (Meller, 1996). By the mid-1980s, the economy was recovering from the effects of the crisis making unemployment declined rapidly from over 10% in 1987 to 7% in 1989, although inflation, at over 20%, was still unsatisfactory. Beginning with a major deterioration in the 1973-74 period, the years from 1974 to 1990 are characterized by larger fluctuations in the concentration of income, several time reaching historically high levels of inequality. In consequence, the military regime left behind a Gini coefficient of household incomes that was 20% higher than in 1973. What happened with the deciles? As it was in previous periods, total inequality seems to be driven by the inequality in the upper part of the distribution. After a vary marked increase in 1974 and 1975, the top decile fluctuated at very high levels during the 1980s reaching its highest value in 1987.

Democracy was re-established in Chile in 1990; because the new authorities accepted the main ingredients of the old model, this shift did not result in any major deviation from the strong market-oriented, export-led growth strategy of the previous regime. However, a reform of the labor code, aimed at boosting the bargaining power of workers, was introduced, and the real minimum wage increased by 28% between 1989 and 1993 (Ffrench-Davis, 2005). As a result of tax reform and higher public-sector revenue generated by seven years of high GDP growth, per capita social expenditure nearly doubled during this period. The minimum wage was again raised in the second half of the 1990s, but, in addition to the emergence of several macroeconomic imbalances during this period, the Chilean economy began to suffer the widespread effects of the Asian crisis. The result was the first GDP contraction since the crisis of the early 1980s.

In a context of solid GDP-growth that characterize the first years of democracy, the Gini coefficient experienced a great decline from 0.56 in 1990 to 0.49 in 1993. In the remainder of the research period, however, inequality followed an inverted-U pattern, reaching its peak at the end of the decade. In contrast to the 1980s, Figure 1 exhibits a remarkable stable inequality bot at the lower and upper part of the distribution during the 1990s.

Behind the pattern of household income inequality described above, within different segments of the labor market the degree of inequality did not followed the same pattern. The distribution of blue-collar hourly earnings (column 2 of Table 1) reached a peak in 1970-73 before trending downwards. White-collar (column 3) and male earnings inequality (column 6) followed the household inequality pattern

Table 1

Average Gini Coefficient of Household Incomes and Earnings by Periods in Great Santiago, 1957-2004

	Households (1)	Blue-Collars (2)	White-Collars (3)	Own-Account (4)	Employers (5)	Males (6)	Females (7)
1957-63	0.474	0.299	0.418	0.525	0.445	0.499	0.547
1964-69	0.496	0.305	0.443	0.527	0.460	0.514	0.539
1970-73	0.475	0.317	0.440	0.510	0.421	0.499	0.522
1974-81	0.525	0.311	0.447	0.542	0.459	0.531	0.504
1982-86	0.562	0.312	0.464	0.525	0.407	0.547	0.500
1987-90	0.580	0.294	0.506	0.559	0.510	0.592	0.544
1991-98	0.522	0.278	0.465	0.521	0.492	0.534	0.478
1999-2004	0.531	0.302	0.491	0.547	0.532	0.553	0.488

Source: Author's calculations from EUSS 1957-2004.

Notes: Calculated using 100 Bootstrap replications. (1) Household income per capita; (2)-(7) Hourly earnings.

more closely, with a long trend upwards from 1970-73 to 1987-90, followed by a slight decline and subsequent increase in the final period. Specially during the 1980s, these changes showed to be highly significance according to the Bootstrap analysis, see Table A1.

The distributive situation of employers and own-account workers, who represent the informal segment of the labor market, showed similar and significant inequality changes in general terms, even though the period from 1982-86 was characterized by a decline, rather than an increase, in the Gini coefficient. Female inequality declined until 1982-86, but followed the pattern of most other groups during the remaining periods. Of these changes, only for the periods 1970-73, 1974-81, 1987-90, and 1991-98 I found evidence of significant changes.

Another aspect of inequality that complements my previous results is the relative average hourly earnings of the different labor market groups analyzed in Table 1. This might give an indication of how the between-group inequality behaved during the research period. The increase in earnings inequality among white-collar workers during the 1980s was accompanied by a growth in the ratio of average white-collar earnings with respect to average blue-collar earnings in the periods 1982-86 and 1987-90. This implies that inequality during the 1980s not only deteriorated within but also between labor market groups. This is even more clearly reflected by employers who's relative earning doubled between the early 1970s and the late 1980s. Another group, however, reports a very different patter. The ratio of male earnings to female earnings, with exception of the 1982-1986 period, declined over the research period. In consequence, the average earnings of male workers were only 30% higher than that of female workers during the 1990s compared with more than 100% in the beginning of the research period.

Table 2

Relative Average Earnings by Periods in Great Santiago, 1957-2004

	White-Collars (1)	Own-Account (2)	Employers (3)	Males (4)
1957-63	2.943	2.154	6.539	2.155
1964-69	2.954	2.073	7.827	1.895
1970-73	2.841	1.929	6.379	1.738
1974-81	2.770	2.204	8.820	1.639
1982-86	3.196	2.070	9.541	1.641
1987-90	3.452	2.281	13.060	1.548
1991-98	2.803	1.982	8.386	1.421
1999-2004	2.782	1.797	8.360	1.324

Source: Author's calculations from EUSS 1957-2004.
Notes: (1)-(3) Relative to the average hourly earnings of blue-collar workers; (4) Relative to the average hourly earnings of female workers.

From my results we can draw the following conclusions: first, the period 1974-91 is characterized by large, positive, and significant household income inequality changes. Second, inequality seems to be driven by the inequality in the upper part of the distribution, while inequality in the lower part played only a marginal role. Third, not only the inequality of household incomes was high during the 1980s, also white-collar and male hourly earnings exhibit a similar deterioration in their distribution. Fourth, while the ratio of average white-collar earnings increased by 25% during the 1980s, it almost doubled for employers.

2.2 Is the Chilean Pattern Different?

In an international context, Chilean inequality is relatively high (Beyer, 1997) and is one of the most unequal in the world (Bravo and Contreras, 1996). One of the latest reports of the World Bank (World Bank, 2003) reveals that, when measuring inequality as the Gini coefficient of the household income per adult equivalent, Chile was the fourth most unequal of 14 Latin American countries in the early 1990s; by the late 1990s, only Brazil had a wider dispersion of household income. Szekely and Hilgert (1999), in a study using data from the mid-1990s, estimated a less extreme position for Chilean inequality–seventh out of 18 Latin American countries. However, if labor income is used as the variable of analysis, Chile moves to the fourth position, and, if only urban areas are analyzed, the country moves into second.

Contrasting the trends reported in Table 1 with the international literature, I found that the deterioration in inequality during the 1980s was not unique to Chile. Gottschalk and Smeeding (1997) suggest that, with the exception of a few countries, almost all industrial economies experienced some increase in wage inequality among prime-aged males during the 1980s. In another work, Gottschalk and Danziger (2005), it is found that both male wage inequality and family income inequality followed a similar pattern during the last decades, increasing during the 1980s, and even during the 1990s but a slower rate. A similar pattern is found by Atkinson (2007) analyzing earnings inequality in U.K. and Germany. Even in some East Asian developing countries such as Hong Kong, South Korea, Singapore, and Taiwan the 1980s was associated with worsening income distribution. A pattern that continued during the 1990s for Hong Kong and Taiwan (Krongkaev, 1994; Zin, 2003):

Studies of Latin America suggest that household inequality increased in six of seven countries for which national data were available for both the beginning and the end of the 1980s. For those countries with only urban data available, three out of six reported higher inequality at the end of that decade (Psacharopoulos *et al.*, 1995). During the 1990s, the picture is less one-sided. The three largest economies in the region report completely different patterns of inequality: a sharp increase in Argentina, remarkable stability in Mexico, and a decline in Brazil. In countries such as Uruguay, Bolivia, Colombia, and Venezuela, inequality increased during the 1990s, but at different rates (World Bank, 2003).

In summary, it is difficult to draw some clear conclusions since almost no country report series of inequality for the whole period covered in this paper. We can, however, say that, on the one hand, the inequality increase observed during the 1980s in Chile was not unique in an international context. On the other hand, the deterioration of inequality in Chile seems to have been more dramatic and widespread than in most other countries.

3 Explaining Income Inequalities

After reviewing the inequality patterns of the last decades, this section examines the results of previous studies on income inequality in Chile. The literature that covers this topic is both vast and varied. One major group of studies uses, as I do in the first part of this article, the EUSS survey. Consequently, these studies have in common the long-run analysis of the data. A second major group of studies relies on the household survey CASEN, which is representative of the entire country, but covers a considerably shorter period of time than EUSS (1987-2003). The majority of these studies employ some type of inequality decomposition, generally by population sub-groups, in order to explain the overall level of inequality. The household is the most common unit of analysis, with household income per capita as the income variable; however, some studies use earnings or wages (male and female separately) as the relevant income variable.

Some results, such as the large percentage of total inequality explained by education, are in line with international research. Others, such as the small share of disparity explained by rural-urban differences, are more unique to Chile. A striking result emerging from these previous studies is that only a few variables, especially education, and, to a lesser extent occupation, are important explanatory factors of inequality in Chile. For instance, while between-group income differences of educational groups explain up to 40% of total disparity (depending on survey, definition of income, year, method, and sample used), other variables, such as family size and household composition (when analyzing household income) or age and sector of employment (when earnings are analyzed), explain only a small percentage of the overall inequality. It is therefore not surprising that the role of education is one of the most studied issues related to inequality in Chile. This is the topic of the next section.

3.1 Education

One important characteristic of the period I analyze is the continued increase in the Chilean's workers level of education, especially in the 1980s and beyond. The average term of education for male workers in the Great Santiago increased from 6.8 in 1958 to 7.9 in 1974, and jumped to over 10 years by 1987 (Contreras, 2002). Especially important to this trend was the growth in the percentage of university-educated individuals, which almost doubled between the mid-1970s and the late 1980s as the result of increased private-sector participation in providing this level of education.

The human capital approach, which assumes that individuals invest in education to maximize the present value of their expected stream of future earnings net of costs, dominates the research on the determinants of labor earnings. Assuming a large time horizon and low educational costs, one can regress log-earnings on years of education and work experience (and its square) to estimate the private return of an additional year of schooling. Thus, an important percentage of the inequality of log-earnings, defined as its variance, could be explained by the inequality of years of schooling and the marginal rate of return to education. It was in this manner that Fields and Yoo (2000) developed an inequality decomposition that makes it possible to analyse the effect of education on inequality. In this decomposition, the share of inequality explained by education is given by the expression:

$$S^{Education}(\ln Y) = \frac{\beta^{Education} * \sigma(Education) * \rho(Education, \ln Y)}{\sigma(\ln Y)} \tag{1}$$

where $Education$ is years of formal education; $\sigma(Education)$ is its standard deviation; $\ln Y$ is log-earning; $\sigma(\ln Y)$ is the standard deviation of $\ln Y$; $\rho(Education, \ln Y)$ is the correlation between $Education$ and $\ln Y$; and $\beta^{Education}$ is the estimated parameter of years of education interpreted as the rate of return to education. Equation 1 indicates that there is a positive relation between the rate of return to education, the inequality of years of education, and the correlation between years of education and log-income; and the percentage of inequality explained by education. On the other hand, the higher the level of inequality, the smaller the percentage explained by education. This methodology was used by Contreras (2002), Contreras (2003), and Amuedo-Dorantes (2005). Their results are summarized in Table 3.

Table 3

Summary of Studies about the Effect of Education on Inequality

Article	Data (1)	Income Variable (2)	Period (3)	Result (4)
Ferreira and Litchfield (1998)[a]	CASEN	Household income/ adult equivalent	1987-94	32%-24%
Contreras (2002)[b]	EUSS	Male wages	1958-96	33-43%
Contreras (2003)[b]	CASEN	Monthly earnings	1990-96	18%-21%
Amuedo-Dorantes (2005)[b]	CASEN	Male wages	1994-2000	13%-14%
		Female wages	1994-2000	18%-16%

Notes: (4) Represent the share of the inequality of respective income variable explained by education. [a] Uses an inequality decomposition of the Theil index. [b] The share of inequality accounted by education is calculated using equation (1).

Contreras's work is the most extensive and informative because it covers several decades. His calculation reveals that among primary-educated workers, the impact of an additional year of education on earnings was relatively constant from 1958 to 1996. Among those with secondary education, the impact of a marginal year of education fell to less than a fifth that of the first years in the study, far below those with only primary education (Table 4). Although the number of those with university education tripled, the return on an additional year of university education went up by 50%, therefore, university education became an increasingly dominant cause of disparity over time, while secondary education's contribution to inequality declined, completely collapsing in the final period.

Table 4

Contribution of Education to the Inequality of Male Wages in Great Santiago, 1958-96

	Total		Return			Share			
	Return	Share	P	S	U	P	S	U	Total
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
1958-65	0.121	38%	0.096	0.168	0.168	27%	45%	28%	100%
1966-70	0.138	39%	0.104	0.134	0.188	26%	35%	40%	100%
1971-75	0.124	34%	0.098	0.106	0.162	27%	30%	42%	100%
1976-80	0.150	42%	0.106	0.122	0.218	24%	30%	46%	100%
1981-85	0.152	38%	0.080	0.132	0.232	18%	31%	51%	100%
1986-90	0.150	39%	0.084	0.074	0.262	19%	18%	63%	100%
1991-96	0.130	32%	0.091	0.033	0.253	22%	8%	70%	100%

Source: Contreras (2002).
Notes: (2) Represents the percentage of inequality explained by education calculated using equation (1); (6)-(8) Represent the percentage of the numbers in column (2) Explained by the different levels of education. P=Primary education, S=Secondary education, U=University education.

Based on the larger sample from the national survey CASEN, Contreras found a 0.10 overall return on education (which is consistent with Arellano and Braun, 1999) when regressing income on gender, experience, level of labor market participation, self-employment, occupation, and education. Education was the most important single factor, accounting for about 20% of the inequality, compared to 5-9% for occupation and self-employment status. Amuedo-Dorantes (2005) also used CASEN data, with separate regressions for male and female workers. Years of schooling was one the most important observable variables, accounting for 13% of wage inequality for males, and slightly more for females.

Ferreira and Litchfield (1998) applied a decomposition of the Theil index using an education partition dividing households into five groups based on the head of household's level of education. The between-group component of this decomposition explained between 24% and 32% of the household income inequality. The second most important variable, occupation, explained no more than a third of the share of inequality explained by education.

There is, then, consistent and convincing evidence of the impact of education on income disparity; the results are particularly pronounced for university education and for males. There are, however, some caveats to be considered. For instance, the direct costs of education are not taken into account in any of the studies surveyed here. The direct cost of education was certainly important during the 1980s, when an increasing proportion of university education was provided by the private sector and tuition fees were introduced in public universities. Moreover, the schooling parameter of the wage equation is biased if unemployment disproportionately affects the less educated; this was clearly the case during the 1980s (Riveros, 1990), when unemployment rose sharply. Accordingly, the results of those studies based on EUSS data should be interpreted cautiously and, in the future, the methods used to analyze the effect of education on inequality should be improved.

3.2 Openness and Trade

The trade liberalization that took place after 1973 is one of the most significant structural changes to the Chilean economy of the last 50 years. Consequently, this shift in policies regarding trade represents one potential driver of income inequality. After the liberalization of the country's external sector, the rates of return on university education were higher than those of any prior period, and, as a result, several groups of the labor market suffered from historically high inequality.

As were many other Latin American countries in the early 1970s, Chile was implementing an ISI strategy. Immediately after the military takeover, the new economic authorities announced a planned shift to a more outward-oriented strategy, thus opening the country to external competition. Import tariffs were reduced in step from an average of 94% in 1973 to a 10% uniform tariff in 1979 (Table 4, below).

After a temporary increase in the early 1980s, the tariff was again gradually reduced, reaching 9.5% in 1990-2000. Exports increased from 9.9% of GDP in 1970-73, of which 80% was copper, to 29.1% of GDP in 1995-2000, of which only 46% was copper (Ffrench-Davis, 2005). Exports increased by 9.5% per year during the 1990s and were a primary source of the high economic growth of this period. As a consequence, the degree of openness of the 1999-200 Chilean economy was nearly double that of 1973.

Table 5

Average Tariff, Openness, Real Exchange-Rate, and Current Account Deficit for Chile, 1973-2000

	Average Tariff (1)	Openness (2)	Real Exchange Rate (3)	Current Account Deficit (4)
1973	94.0	29.6	65.1	n.a.
1974-79	35.3	45.6	73.2	4.3
1980-82	10.1	44.3	57.6	10.2
1983-85	22.7	47.9	79.1	8.3
1986-89	17.6	57.5	106.6	3.5
1990-95	12.2	57.1	99.5	2.5
1996-98	11.0	56.1	80.3	5.4
1999-2000	9.5	59.4	84.1	0.8

Sources: (1) and (3) Ffrench-Davis (2005: 168); (2) Heston *et al.*; (4) Beyer *et al.* (1999) and Corbo and Tessada (2003); n.a. = not available.

Notes: (1) in percentage; (2) (Exports+Imports)/GDP; (3) 1986=100; (4) as percentage of GDP.

The Heckscher-Ohlin Stolper-Samuelson (HOSS) framework, in which the relevant factors of production are skilled and unskilled workers, has long been the dominant approach to analyze the effect of trade on inequality. Trade is a substitute for factor mobility in equalizing relative factor incomes across trading nations, first by equalizing relative commodity prices, and thereafter wages. The Stolper-Samuelson theorem states that an exogenous reduction in the relative price of a commodity, for instance through a reduction of tariffs, reduces the rate of return of the factor used intensively in the production of that commodity (in a 2-commodity model), and increases the rate of return of the other factor. Under this theorem, then, increased openness and trade in a country such as Chile would shift production towards sectors that use unskilled labor more intensively, thus increasing the demand for, and wages of, unskilled workers and reducing overall income inequality.

Beyer *et al.* (1999) analyzed the direct effect of openness on inequality in Chile. First, an earnings regression on EUSS data was estimated for each year from 1960 to 1996. Then, using time-series techniques, the differences between returns on university and returns on primary education ($DCG - DEG$) were regressed on three different factors: the sum of imports and exports as a share of GDP ($Openness$); the index of the producer price of textile goods (P_{Tex}), representing the price of unskilled- labor intensive goods; and the proportion of university-educated workers ($University$), representing the supply of skilled workers. The series used were integrated of order 1, as well as co-integrated. The estimation of the regression generated the following results:

$$(DCG - DEG) = \underset{(8.81)}{1.908} + \underset{(3.64)}{0.0131} * Openess - \underset{(\text{-}2.09)}{0.357} * P_{Tex} - \underset{(\text{-}2.61)}{0.027} * University \quad (2)$$

with t-values in parenthesis; $R^2 = 0.610$; $D.W. = 1.419$; and $ADF = -4.924$.

In line with the Stolper-Samuelson theorem, the statistically significant negative parameter of P_{Tex} indicates that a lower price on goods intensive in unskilled labor increases the difference in the return on education. This result is also consistent with the increase in the rate of return on university education relative to primary education witnessed in the second half of the 1970s and mid-1980s. This difference increased from 0.064 in the early 1970s to 0.178 in the late 1980s. The one-third tariff reductions of the late 1970s and late 1980s reduced the price of imported goods, including textile goods, and thereby reduced the price of commodities and unskilled wages; the result was a widened gap between rates of return on education. *Openness* has a statistically significant positive effect; that is, the higher the degree of openness to international trade, the greater the difference between returns on university and primary education. Based on the distributional predictions of the simple version of the HOSS framework, this result is not expected. Admittedly, the changing pattern of Chilean exports was consistent with the predictions of HOSS because trade liberalization implied that the sectors in which Chile has comparative advantages, such as agriculture, forestry, and fishing, increased their exports. Moreover, unskilled-labor intensive industries, such as shoes and leather goods, were important components of the increase in industrial exports. Ultimately, however, the distributional effect of the trade liberalization was a higher level of inequality (and not a lower one, as suggested by trade theory).

Robbins (1994) used EUSS data and Katz and Murphy's (1992) methodology to study shifts in supply and demand for different gender and educational groups in the period 1966-92. For the period of 1975-90, between-sector demand changes favored male workers with secondary or university education, and female workers with university or special education. Within-sector demand changes for the same period favored more educated workers to an even greater degree. Only from 1991-92 did demand shifts favor primary education. A potential weakness of this study is that it covers only the Greater Santiago area, where important sectors such as agriculture and mining are absent. Therefore, Bravo and Contreras (1999) applied the same method, but instead used the CASEN data for the period of 1990-96. They concluded that between-sector demand changes were positive in some cases and negative in others, but in general were very small. Within-sectors, demand shifts were negative and small in most cases. One important exception was the group of male university-educated workers, who experienced a positive and high within-sector demand shift. Thus, most of the results from available studies indicate that openness was followed by an increased demand for highly educated workers in Chile. But why was this the case?

Pavcnik (2003) offers one potential explanation. She suggests that investment was the link between trade liberalization and increased demand for skilled workers in Chile. Trade liberalization reduced the relative price of imported machinery, materials, and technology, and invited increased competition from imported products. In this environment, manufacturing plants increased investment, thereby increasing the demand for skilled labor, a complement of the plants. Pavcnik also suggests that the use of imported materials, foreign technical assistance, and foreign patent technology (all proxies for foreign technology) by manufacturing plants was not associated with the increased demand for skilled labor. Underlying this result was that only certain plants within the different sectors adopted such foreign technology, and the majority of those plants employed relatively more skilled labor even before they imported foreign technology.

Beyer *et al.* (1999) offer a second possible explanation for the spike in demand for educated workers: the increased openness in Chile induced a more intense exploitation of natural resources. Although the share of natural-resource based goods has remained at roughly 80% of total exports for many years, the number of exported products has increased significantly over the same period. If the sectors producing these goods induced a demand shift biased in favor of skilled or highly educated workers, it may explain the positive effect of openness on the widening of the wage structure. However, according to Wood (1997), there is no empirical evidence that the primary production or primary processing sectors in Latin American countries are skill intensive. A third possible explanation of the pattern of inequality during Chile's trade liberalization is a concurrent liberalization of the labor market. According to Wood (Wood, 1997), rejecting the effect of the reduction in union power during this period, as some authors do, is not totally convincing. Other scholars have pointed out that "The combination of an open economy and a flexible labor market is believed to be the cause of many growing socioeconomic ills, including income inequality"(Gill and Montenegro, 2002).

3.3 Labor Market Aspects

Given that the data reveal high levels of income inequality during periods of labor-rights suppression, changing labor market policies have also been posited as a potential source of Chile's increasing income disparity since the 1980s. Edwards and Edwards (2000) identified four main periods of labor market policies during the last decades: 1966-73, 1974-79, 1980-90, and 1991 onwards. They described the 1980s, characterized by the removal of collective bargaining and revocation of the control of economic authorities to adjust wages, as the least restricted period.

Small average wage growth and even a decline in the minimum wage are also distinguishable features of Chile in the 1980s (Table 6). Additionally, the average unemployment rate for the period was the highest of the last four decades. Unfortunately, there is no available work on the effect of changes in collective bargaining on inequality, but there are, however, several studies that account for the effect of minimum wage, female participation, and unemployment. The main results of these studies are summarized in Table 7.

Table 6

Labor Market Indicators for Chile, 1970-2000

(All numbers represent %)

	Collective Bargaining (1)	Real Average Wage Growth (per year) (2)	Real Minimum Wage Growth (per year) (3)	Female Participation (4)	Unemployment Rate (5)
1970-73	26.1	-5.1	35.3	33.6	4.7
1974-79	0.0	0.9	-9.2	32.6	13.7
1980-90	13.4	1.3	-0.7	34.6	14.7
1991-2000[a]	16.7	3.6	5.7	39.9	8.5

Sources: (1)-(3) Cortázar (1997) and Ffrench-Davis (2005); (4) Larrañaga (2001); (5) Chile Social and Economic Indicators 1960-2000.

Notes: (1) Percentage of wage earners covered by collective agreements; (4) The values of the last three periods represent 1974-81, 1982-86, and 1991-98; [a] Refers only to the period 1991-93 for collective bargaining.

Larrañaga (2001) used a regression approach with the Gini coefficient of per capita household income as the dependent variable, and unemployment and female participation as the explanatory variables. He found only a small impact of unemployment on inequality: 0.039 points (7.5%) of the average Gini coefficient for these years (0.517). Meller *et al.* (1996) instead used the share of total income of the different quintile groups as the dependent variable, ultimately showing that unemployment had a significant positive impact on inequality because higher unemployment increases the income share of the top quintile at the expense of the bottom four. When analyzing the impact of minimum wage on inequality, Meller *et al.* (1996) found a significant negative impact, but only in the first and second quintiles.

Table 7

Summary of Studies about Labor Market and Inequality

Article	Data (1)	Income Variable (2)	Period (3)	Result (4)
		Unemployment		
Larrañaga (2001)	EUSS	Household income/capita	1957-97	positive effect
Meller (1996)	EUSS	Household income/capita	1959-96	positive effect
		Minimum wage		
Meller (1996)	EUSS	Household income/capita	1959-96	negative effect
		Female labor force participation		
Larrañaga (2001)	EUSS	Household income/capita	1957-2001	positive effect
		Informal employment		
Uthoff (1986)	EUSS	Monthly earnings	1969, 1978	positive effect
Amuedo-Dorantes (2005)	CASEN	Male wages	1994-2000	2%-3%
		Female wages	1994-2000	1%-2%
		Occupation		
Ferreira and Litchfield (1998)	CASEN	Household income/ adult equivalent	1987-94	10%-8%
Amuedo-Dorantes (2005)	CASEN	Male wages	1994-2000	17%-21%
		Female wages	1994-2000	20%-24%

Notes: (4) A positive effect means that a higher rate of unemployment induces a higher level of inequality. A negative effect means that a higher minimum wage induces a lower level of inequality. The numbers represent the percentage of inequality explained by respective variable.

Considering that most of the workers affected by the minimum wage are found in the lowest income groups, this result is not surprising.

The expansion of female participation in the labor force, as occurred after 1974-79, had two potential opposing effects on inequality. Because female workers are, on average, paid less than are males, and because they frequently work part-time, their increased participation might place more workers at the bottom of the earnings distribution. On the other hand, assuming a positive correlation between spouses' levels of education, increased labor force participation for women implies a disproportionate advantage for high-earning, highly educated households. Larrañaga (2001) found a positive relationship between this variable and inequality. An increased female labor force explained 0.166 points of the average Gini coefficient (32%), which is substantially more than was explained by unemployment. This most probably indicates that the second effect was stronger, and that females from households with higher incomes increased their participation in the labor market.

Occupation is the single most important factor in the realm of labor markets. It is the second most important factor in explaining the inequality of household incomes and, in the work of Amuedo-Dorantes, which analyzed the distribution of male and female wages, it is even more important than education. Armuedo-Dorantes focused her study on the effect of Chile's growing informal employment on inequality. She defined the informal sector as wage and salary workers without contract, and concluded that, from 1990 to 2000, informal wage employment increased from 10% to 18% for males, and from 12% to 26% for females. Because informal wage employment is characterized by a narrower dispersion of and lower average earnings, its increased incidence should affect the overall distribution of earnings. However, at 1-3%, the contribution of informal employment to the explanation of wage-rate inequality during the 1990s is modest. Informal employment was, however, more important to explain the change, rather than the level, of inequality in this period.

Although different aspects of the labor market has bee widely discussed in an inequality context, some issues have not yet been explored. There are few studies of income inequality in the labor market, using CASEN data, in which multiple disaggregated groups are analyzed. Males and females in the labor force are separately analyzed in several existing studies, but the self-employed have been largely neglected in the Chilean literature. Due to its increased importance to the labor markets of many countries, this class of worker has begun to attract increased attention in international literature (Le, 1999), but no inequality analysis of them exists in the Chilean case.

3.4 Public Policies

Social policies were important to the democratic government that came to power in Chile in 1990. The new leadership proved willing to introduce reform and direct increased tax revenues to actively reduce poverty and inequality without jeopardizing the stability or growth of the economy. This strategy was deemed 'growth with equity' (*crecimiento con equidad*) and required not only a new structure of social expenditures but also new governmental institutions and a tax reform. Before the tax reform of 1990, almost 50% of tax revenue was generated by the VAT, while only 18% came from income taxes. The tax rate on corporate profits was among the lowest in the world (10%) (Marcel, 1997). The reform increased the VAT from 16% to 18% and the income tax to 15%. These changes generated an additional 800 million of dollars of tax revenue each year; total tax revenue as a percentage of GDP increased from 14% in 1990 to over 16% in 2000. Even though the extra revenue from these measures was important to increasing public outlays, 69% of the increased tax collection of the 1990s was the result of growth, while only 31% was due to the tax reform itself (Arellano, 2004). The focus on social policies in this period was also seen in the creation of new governmental institutions catering to some of the most vulnerable groups of the Chilean population, including: FOSIS[2], SERNAM[3], INJ[4], CONADI[5] and FONADIS[6]. Social policies become also more strongly targeted at poorest households during the 1990s. Each new increase in resources and each new program were focused at the old with lowest pension, or at poorest schools, or households with greatest needs. Public subsidies, for instance, were redistributed to the first two quintiles, which received 57% of the total of monetary subsidies in 1990 but increased to 73% in 2000.

Bravo *et al.* (2002) employ a very ambitious approach to evaluate the distributive impact of relevant public policies. Most income inequality studies rely on different definitions of income and simulate the distributive impact of public policies using quintile data. However, none takes into account the effect of both monetary and in-kind transfers at the household level. Non-monetary transfers such as subsidies for education, health, and housing, free up income for consumption. Therefore, including them in total household income generates a measure of income that is more closely related to consumption and welfare.[7] To estimate the monetary value of different in-kind subsidies, Bravo *et al.* used detailed information from different governmental

[2]Fondo de Solidaridad e Inversion Social

[3]Servicio Nacional de la Mujer

[4]Instituto Nacional de la Juventud

[5]National Corporation for Indigenous Development

[6]Fondo Nacional de la Discapacidad

[7]Besides the different definitions of income included in the CASEN survey, Bravo *et al.* introduce the definition of *net income of social policies*, which reduces from total income monetary and housing subsidies. *Income with social policies* is defined as household net income per capita plus monetary transfers and plus the value of in-kind subsidies.

institutions. The monetary value of each in-kind transfer was added to the income of the household that benefited from that transfer. Finally, income inequality measures were calculated before and after the transfers. An important assumption of the study is that one monetary unit of in-kind subsidy is equivalent to one monetary unit of available income for the beneficiary [8]

Table 7 shows that social policies have an important inequality-reducing effect. The impact of monetary transfers on the Gini coefficient is relatively small (column 2), whereas the total impact of social policies on both the Gini coefficient and the ratio is remarkable (column 3).

Table 8

Inequality Indicators with and without the Effect of Social Policies, CASEN 1998

Indicator	Income Net of Social Policies (1)	Total Household Income (2)	Income with Social Policies (3)
Share of first quintile	3.06	3.43	5.16
Share of second quintile	6.68	6.94	8.20
Share of third quintile	10.81	10.95	11.60
Share of fourth quintile	18.31	18.29	18.02
Share of fifth quintile	61.14	60.39	57.02
Total	100.00	100.00	100.00
Ratio fifth to first	20.00	17.60	11.10
Gini Coefficient	0.564	0.554	0.503

Source: Bravo *et al.* (2002).

Notes: (1) Net income per capita: per capita income of the household minus monetary subsidies and housing. (2) Per capita income of the household including autonomous income, monetary subsidies and rent attributed to the house. (3) Per capita income with social policies: net income per capita plus monetary and in kind subsidies (housing, health and education).

[8] Relaxing this assumption, they found that even when 20-30% of each monetary unit of public expenditure could be transformed into consumption, a substantial reduction in inequality is achieved.

While monetary transfers reduce the Gini coefficient from 0.564 to 0.554, social policies reduce it to 0.503. The largest effect of social policy is found in the ratio between the income shares of the fifth quintile and first quintiles: when net income of social policies is used, the ratio is cut nearly in half. Education is the major source of the inequality-reducing power of social policies. On its own, education reduced the Gini coefficient to 0.529 and the ratio to 14, and accounted for almost 45% of the total reduction in inequality achieved by social policies in aggregate. Because public policies are an important factor in reducing inequality, it is clear that calculations based on only monetary incomes can distort the measurement of income disparity. This becomes even clearer when making international comparisons, because public policies vary widely across countries.

4 Conclusions

This paper described the evolution of the inequality of household income and earnings across different groups in the Chilean labor market since the late 1950s. It also summarized the existing explanations of the level of and changes in this inequality. Except for the years of 1970-73, average household income inequality increased by period from 1957-63 to 1987-90, and declined in the 1991-98 period. The inequality of male and white-collar hourly earnings and the rate of return on university education have all generally followed the pattern of overall income inequality.

Thus, although inequality was lower in the 1990s than in the previous decade, it remained high; additionally, the distributive situation of earnings within several specific groups of the labor market was still unsatisfactory. These facts are puzzling in light of the knowledge that during the 1990s the Chilean economy experienced three major inequality-decreasing events: rapid GDP growth, an increasing minimum wage, and a relatively low rate of unemployment (Larrañaga, 2001; Meller *et al.*, 1996). One possibility is that the 1990s expansion of the economy, based on an export-oriented and natural resource-based strategy, did not have an effect on the labor market sufficient to generate a substantial impact on aggregate measures of household income and earnings inequality. Alternatively, because inequality-increasing factors also emerged during this period, only small overall inequality changes resulted at the aggregate level. One important such factor might have been the changing structure of the labor market. The informal sector grew during the 1990s (Amuedo-Dorantes, 2005); if low-paid and low-regulated informal employment grew at the expenses of formal employment, this might have widened the earnings distribution. Moreover, since informal employment provides more flexible working schedules, a surge in this sector may affect the dispersion of earnings through an increase in the inequality of working hours and thereby of monthly earnings. Also, a shift towards professional and away from unskilled occupations occurred during the 1990s. Taking these two findings together, workers in the upper as well as in the lower part of the wage distribution increased their share of total Chilean employment during the 1990s. Overall, this should have widened the distribution of earnings.

There is also evidence that the pattern of inequality in the upper part of the distribution drives general earnings inequality. In this context, quintile regressions are used to analyze whether the determinants of inequality among high-earning workers differ from the determinants among their low-earning counterparts. Although this approach has begun to appear more frequently in the international literature on inequality, it is still relatively unexplored in Chile. It would be informative to investigate whether some of the results of the studies approached here, such as the important role of education and occupational structure, apply equally to the upper and lower ends of the distribution, and whether GDP growth has resulted in equal earnings growth rates across the entire earnings distribution.

Another area that deserves future attention is the analysis of the effect of data contamination on the inequality measures obtained from CASEN. CASEN is perhaps the best national representative data to analyze income inequality in Chile. However, it covers a considerably shorter period of time than does EUSS (every two years from 1987), its classification of occupations and industrial sectors changed in 1992, and several corrections have been made with regard to underreporting and missing income variables. Unfortunately, there is only limited information on how missing values are distributed, and no studies have yet analyzed the effect of the corrections on income inequality indicators; however, several methods to correct for missing values, other than the mean value imputation currently applied to CASEN, are now available. Moreover, almost no studies on the statistical significance of inequality indicators using CASEN data exist. As I mention in section two of this paper, the use of the Bootstrap technique is relatively uncomplicated and, at a low computational cost, might generate important insights into Chile's income inequality during the high-growth 1990s. Several additional avenues of investigation will be pursued in the future: the effect of the changing structure in the labor market, differences in the determinants of inequality in the different parts of the distribution, and the significance of income inequality changes during the 1990s.

References

Amuedo-Dorantes, C. (2005): "Work Contracts and Earnings Inequality: The Case of Chile," *Journal of Development Studies*, 41:4, pp. 589-616.

Arellano, J. P. (2004): "Políticas Sociales para el Crecimiento con Equidad, Chile 1990-2002," *Serie Estudios Socio/Económicos*, 26.

Arellano, M. S., and M. Braun (1999): "Rentabilidad de la Educación Formal en Chile," *Cuadernos de Economía*, 107, pp. 685-724.

Atkinson, A. B. (2007): "The Long Run Earnings Distribution in Five Countries: "Remarkable Stability", U, V or W," *Review of Income and Wealth*, 53(1), pp. 1-24.

Beyer, H. (1997): "Distribución del Ingreso: Antecedentes para la Discusión," *Estudios Públicos*, 65 (Verano), pp. 1-54.

Beyer, H., P. Rojas, and R. Vergara (2000): "Apertura Comercial y Desigualdad Salarial en Chile," *Estudios Públicos*, 77 (Verano), pp. 70-95.

Bravo, D., and D. Contreras (1999): "La Distribución del Ingreso en Chile 1990-1996: Análisis del Impacto del Mercado del Trabajo y las Políticas Sociales," Universidad de Chile, Mimeo.

Bravo, D., D. Contreras, and I. Millán (2002): "The Distributional Impact of Social Expenditure: Chile 1990-1998," Background Paper 2 to *Chile's High Growth Economy, Poverty and Income Distribution 1987-1998*, World Bank Country Study.

Chile Social and Economic Indicators 1960-2000 (2002): *Banco Central de Chile*, Santiago, Chile.

Contreras, D. (2002): "Explaining Wage Inequality in Chile: Does Education Really Matter?" *The Journal of Income Distribution*, 11:1, pp. 53-76.

Contreras, D. (2003): "Poverty and Inequality in a Rapid Growth Economy: Chile 1990-1996," *The Journal of Development Studies*, 39:3, pp. 181-200.

Corbo, V. (1993): "Economic Reforms in Chile: an Overview," Universidad Católica de Chile Working Paper 160.

Corbo, V., and J. A. Tessada (2003): Growth and Adjustment in Chile: A Look at the 1990s, Banco Central de Chile Working Paper 2004.

Cortázar, R. (1997): "Chile: The Evolution and Reform of the Labor Market," Chapter 8 in *Labor Markets in Latin America: Combining Social Protection with Market Flexibility*, S. Edwards and N. C. Lustig (eds.), Brookings Institution Press.

Edwards, S., and A. C. Edwards (2000): "Economic Reforms and Labour Markets: Policy Issues and Lessons from Chile," *Economic Policy*, 15:30, pp. 181-230.

Ferreira, F. H. G., and J. Litchfield (1998): "Calm After the Storms: Income Distribution in Chile, 1987-94," Policy Research Working Paper 1960, World Bank.

Fields, G. (2001): *Distribution and Development: A New Look at the Developing World* , MIT Press.

Fields, G., and G. Yoo (2000): "Falling Labor Income Inequality in Korea's Economic Growth: Patterns and Underlying Causes," *Review of Income and Wealth*, 46:2, pp. 139-160.

Ffrench-Davis, R. (2005): *Economic Reforms in Chile, From Dictatorship to Democracy.* Michigan: The University of Michigan Press.

Gill, I. S., and C. E. Montenegro (2002): "Responding to Earnings Differentials in Chile," Chapter 6 in *Crafting Labor Policy: Techniques and Lessons from Latin America*, I. S. Gill, C. E. Montenegro, and D. Dömeland (eds.), World Bank and Oxford University Press.

Gottschalk, P., and S. Danziger (2005): "Inequality of Wage Rates, Earnings and Family Income in the United States, 1975-2002," *Review of Income and Wealth*, 51(2), pp. 231-254.

Gottschalk, P., and T. M. Smeeding (1997): "Cross-National Comparisons of Earnings and Income Inequality," *Journal of Economic Literature*, Vol. XXXV, pp. 633-687.

Heston, R. S., R. Summers, and B. Aten (2006): Penn World Table Version 6.2, *Center for International Economic Comparisons of Production, Income and Prices at the University of Pennsylvania*.

Katz, L., and K. M. Murphy (1992): "Changes in Relative Wages, 1963-1987: Supply and Demand Factors," *The Quarterly Journal of Economics*, 107:1, pp. 35-78.

Krongkaew, M. (1994): "Income Distribution in East Asian Developing Countries: An Uppdate," *Asian-Pacific Economic Literature*, 8:2, pp. 58-73.

Larraín, F., and P. Meller (1991): "The Socialist-populist Chilean Experience: 1970-73," pp. 175-222 in *Macroeconomic of Populism in Latin America*, R. Dornbusch and S. Edwards, (eds.), National Bureau of Economic Research and University of Chicago Press.

Larrañaga, O. (2001): "Distribución de Ingresos en Chile: 1958-2001, Department of Economics," Universidad de Chile, Working Paper 178.

Le, A. T. (1999): "Empirical Studies of Self-employment," *Journal of Economic Surveys*, 13:4, pp. 381-416.

Litchfield, J. (2002): "Updated Income Distribution and Poverty Measures for Chile 1987-1998," Background Paper 1 to *Chile's High Growth Economy, Poverty and Income Distribution 1987-1998*, World Bank Country Study.

Meller, P. (1996): *Un Siglo de Economía Política Chilena (1890-1990)*, Editorial Andres Bello.

Meller, P., R. O'Ryan, and A. Solimano (1996): "Growth, Equity, and the Environment en Chile: Issues and Evidence," *World Development*, 24:2, pp. 255-272.

Marcel, M. (1997): "Póliticas Públicas en Democracia: El Caso de la Reforma Tributaria de 1990 en Chile," *Coleción Estudios CIEPLAN*, 45, pp. 33-83.

Mills, J. A., and S. Zandvakili (1997): "Statistical Inference Via Bootstrapping for Measures of Inequality," *Journal of Applied Econometrics*, 12, pp. 133-150.

Pavcnik, N. (2003): "What Explains Skill Upgrading in Less Developed Countries?" *Journal Of Development Economics*, 71:2, pp. 311-328.

Palma, A. (2006): "Price and Composition Effects on the Rise and Fall of Wages Inequality in Santiago" *Working Paper*, Department of Economics, Göteborg University.

Psacharopoulos, G., S. Morley, A. Fiszbein, H. Lee, and W. Wood (1995): "Poverty and Income Inequality in Latin America During the 1980s," *Review of Income and Wealth*, 41:3, pp. 245-264.

Riveros, L. A. (1990): "The Economic Return to Schooling in Chile: An Analysis of its Long-term Fluctuations," *Economics of Education Review*, 9:2, pp. 111-121.

Robbins, D. (1994): "Relative Wage Structure in Chile, 1957-1992: Changes in the Structure of Demand for Schooling," *Estudios de Economía*, 21, pp. 51-78.

Szekely, M., and M. Hilgert (1999): "What is Behind the Inequality We Measure? An investigation Using Latin American Data," *Luxembourg Income Study Working Paper No. 234*.

Wood, A. (1997): "Openness and Wage Inequality in Developing Countries: The Latin American Challenge to the East Asian Wisdom," *World Bank Economic Review*, 11:1, pp. 33-58.

World Bank (1997): "Chile and Poverty and Income Distribution in a High Growth Economy: 1987-1995," World Bank Document 16377-CH.

World Bank (2003): Inequality in Latin America and the Caribbean: Breaking with History?

Zin, R. H. M. (2005): "Income Distribution in East Asian Developing Countries: Recent Trends," *Asian-Pacific Economic Literature*, pp. 36-54.

Table A1

Significance of the Changes of Average Gini Coefficients of Household Incomes and Earnings by Periods in Great Santiago, 1957-2004

		Households (1)	Blue-Collars (2)	White-Collars (3)	Own-Account (4)	Employers (5)	Males (6)	Females (7)
1964-69	$\Delta Gini$	0.022*	0.006*	0.026*	0.004	0.017	0.014*	-0.007
	(95% CI)	(0.017;0.027)	(0.000;0.012)	(0.016;0.038)	(-0.001;0.016)	(-0.017;0.049)	(0.006;0.023)	(-0.016;0.002)
1970-73	$\Delta Gini$	-0.021*	0.013*	-0.004	-0.017*	-0.037*	-0.013*	-0.016*
	(95% CI)	(-0.026;-0.017)	(0.007;0.019)	(-0.012;0.006)	(-0.027;-0.006)	(-0.063;-0.010)	(-0.020;-0.005)	(-0.025;-0.008)
1974-81	$\Delta Gini$	0.051*	-0.006	0.008	0.032*	0.034*	0.031*	-0.020*
	(95% CI)	(0.046;0.055)	(-0.014;0.000)	(-0.001;0.015)	(0.022;0.041)	(0.007;0.066)	(0.024;0.039)	(-0.028;-0.013)
1982-86	$\Delta Gini$	0.036*	0.001	0.016*	-0.016*	-0.048*	0.016*	-0.004
	(95% CI)	(0.031;0.042)	(-0.006;0.008)	(0.011;0.023)	(-0.029;-0.005)	(-0.071;-0.024)	(0.010;0.022)	(-0.012;0.003)
1987-90	$\Delta Gini$	0.018*	-0.019*	0.043*	0.035*	0.099*	0.044*	0.044*
	(95% CI)	(0.001;0.026)	(-0.026;-0.012)	(0.034;0.051)	(0.020;0.052)	(0.073;0.133)	(0.037;0.054)	(0.033;0.058)
1991-98	$\Delta Gini$	-0.058*	-0.016*	-0.041*	-0.039*	-0.014	-0.057*	-0.065*
	(95% CI)	(-0.066;-0.050)	(-0.021;-0.011)	(-0.049;-0.033)	(-0.055;-0.022)	(-0.040;0.008)	(-0.065;-0.049)	(-0.077;-0.054)
1999-2004	$\Delta Gini$	0.009*	0.024*	0.026*	0.031*	0.036*	0.017*	0.009
	(95% CI)	(0.003;0.015)	(0.0178;0.030)	(0.018;0.034)	(0.018;0.045)	(0.012;0.062)	(0.010;0.026)	(-0.004;0.019)

Source: Author's calculations from EUSS 1957-2004.

Notes: Calculated using 100 Bootstrap replications. * Represents that the inequality change was significant.

Size, Significance, and Sources of Recent Income Inequality Changes in Chile

ALEXIS PALMA †‡

Abstract

This paper analyzes the size, significance, and sources of income inequality changes in Chile during the period 1990-2003. In general, the results indicate that Chile's distribution of income deteriorated at the end of the decade compared with the early or the mid-1990s.

Education was found to be the single most important source of inequality, accounting for 20%-40% of total wage and salary and self-employment income inequality and over 60% of the inequality change between 1996 and 2003. For the first of these income variables this pattern was in turn driven by a strong increase in the wage and salary of highly educated workers in primary and industry sectors and by an increased population share of the highly educated employees in service sectors. Underlying the large, positive, and significant between-education component of the inequality change of self-employment incomes was the contribution of highly educated in service sectors through an increase of their income over most of the 1990s.

Analyzing the household income inequality by income source indicates that the self-employment income and the wage and salary of university educated individuals accounted for over 26% of the household income inequality in 1990. In 2003 this share had went up to near 40%.

†Department of Economics, Göteborg University, Box 640, SE-405 30 Göteborg, Sweden. Phone: +46 (0)31 786 13 67. E-mail: alexis.palma@economics.gu.se

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1 Introduction

The question of what determines the level of income inequality in a given society has been a widely debated issue for many years. However, the interest on this topic has fluctuated from period to period, having lately not only been *back in from the cold* (Atkinson, 1997), it has become a *hot issue of research* (Bourguignon, 2001). At the same time, thanks to the increased supply of high quality household data from developing countries, the research in this subject has expanded to include the study of income distribution in these economies (Kanbur and Lustig, 1999).

Chile is a developing country where national representative high quality data have increased substantially since the late 1980s with the collection of the household survey *Caracterización Socioeconómica Nacional* (CASEN). Other relevant data do exist but CASEN is probably the most ambitious effort made by Chilean authorities to obtain detailed information on income and other socioeconomic variables of Chilean households, considering the large number of households surveyed, the extensive number of variables included, and its national representativeness. As a direct result of the survey, research on income inequality has expanded, and, consequently, several important related insights have emerged during the 1990s.

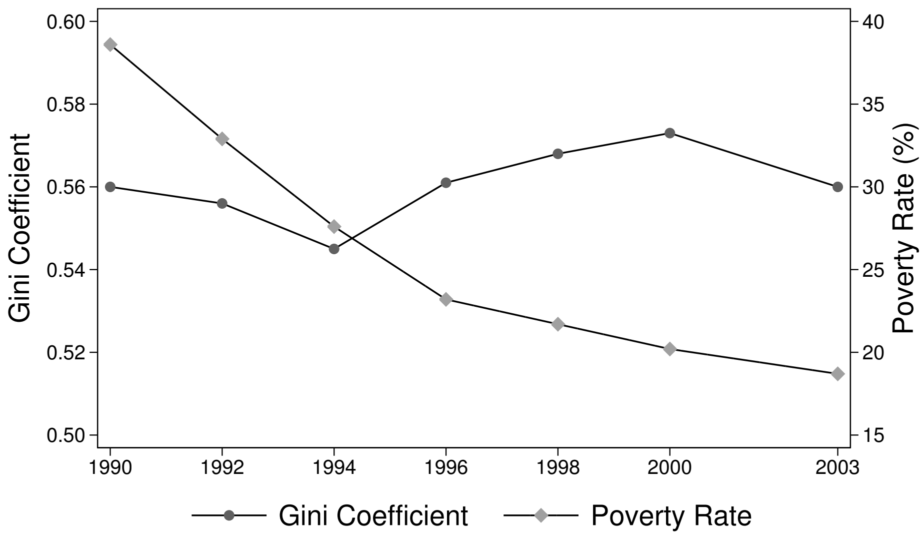
First, although Chile has been rather successful in its attempt to reduce poverty, the country has failed to reduce its high concentration of income.[1] One of Chile's most important economic achievement during the 1990s was perhaps its rapid reduction of poverty, which declined by more than 50% between 1990 and 2003 (Figure 1). Inequality, defined as the Gini coefficient of household incomes, on the other hand, declined only slightly in 1992 and 1994, and actually increased in the

[1]See Palma (2005) for a survey of studies on income inequality in Chile conducted since the early 1990s.

period from 1996 to 2000; as a result, Chile ended the decade with a slightly more unequal income distribution than that with which it began.

Figure 1

Gini Coefficient of Household Income per Capita and Household Poverty Rate, 1990-2003



Source: Author's calculation from CASEN.

Notes: The poverty rate is calculated using the variable *Poverty line* in CASEN which takes the value 1 if the individual is poor. The income variable used to calculate inequality is the total monetary income of the household divided by the number of household members.

Second, all available research indicates that education is by far the most important factor in explaining the dispersion of most income variables in Chile. When analyzing household incomes, level of education accounts for 20%-30% of total existent inequality. Also, when analyzing monthly earnings, education explains 18%-21% of the measured disparity (Ferreira and Litchfield, 1998; Contreras, 2003). Third, and also important to explaining Chile's income inequality, although to a lesser degree than education, are the income differences between different types of occupations.

In fact, between-occupation inequality accounts for 8%-10% of household income inequality, 17%-21% of wage inequality for males, and as much as 20%-24% of female wage inequality (Ferreira and Litchfield, 1998; Amuedo-Dorantes, 2005). Fourth, public policies in the 1990s reduced the Gini coefficient of household incomes from 0.554 to 0.503; because they were largely targeted at the poorest households of Chilean society, these policies increased the income share of these households and reduced inequality (Bravo *et al.*, 2002).

Despite these findings, several issues remain to be investigated in an income inequality context. One such issue is whether education and occupation continue to play an important role in the explanation of Chilean inequality and what part, if any, they have had in the worsened distribution of incomes after 1994.

Another issue in need of further study is the statistical significance of income inequality changes. Although Contreras (1996) conducted such an analysis, his work covered only the period 1987-92. The lack of significance analyses in income inequality studies is not unique to Chile. For a long period of time, point estimates of inequality accompanied by precision measures, such as standard errors, are the exception rather than the rule. And while precision measures should be included in every study, the need for significance analysis is even greater when changes in the concentration of income are relatively small, as they were in Chile during the 1990s.

Additionally, there is a lack of information regarding the distributive situation in different segments of the labor market. The few exceptions look separately at male and female workers, but offer no other labor partitions such as wage and salary workers, and the self-employed. This analysis is important because self-employed workers comprise nearly 25% of the occupied labor force and are the focus of increasing international attention (Le, 1999; Parker, 1999); despite these facts, the self-employed remain largely neglected in exiting Chilean literature.

The present paper addresses these important, but unexplored issues of Chilean income inequality. I independently analyze the inequality of wages and salaries and of self-employment incomes. In this context, the paper aims to answer the following questions: Is there any difference between the socioeconomic composition of wage and salary workers and the self-employed? Is there a difference between the level of inequality among wage and salary workers and the disparity among the self-employed? Is the inequality composition (within-group, between-group) different in these two groups? The inequality of household incomes is then analyzed by income sources to identify how wages and salaries, self-employment incomes, and other income sources affect the distribution of income.

The results reveal that both household income inequality and wage and salary inequality deteriorated significantly in 2000 compared with 1994. For self-employment incomes, the significant changes occurred in 1998 compared to 1990, when inequality declined, and in 2000-03 compared to 1996-98, when inequality increased.

Education, accounting for 20%-40% of the level of inequality, is the single most important explanatory variable for wage and salary and self-employment income inequality, followed by occupation which explains between 15% and 20%. Despite that no significant change in inequality were found for wages and salaries comparing 1996 with 1990 and 2003 with 1996, the between-education component of the inequality changes for these years were large, positive, and significant. This is also true for self-employment incomes.

The decomposition of the household income inequality by income source indicates that the importance of the wage and salary of workers with university education increased markedly over the research period. Together with the earnings of their self-employed counterparts, these two income sources accounted for over 40% of the total inequality. This share represent a 50% increase compared with their explanatory power in 1990.

The paper is organized as follows. The next section presents the inequality indicators used to analyze the level and change of inequality. The third section presents the main characteristics of the CASEN survey and the definition of income variables. The fourth section presents an overview of macroeconomic and socioeconomic data to provide a background picture of the Chilean economy during the 1990s. Following these sections, I present income inequality indicators calculated using wages and salaries, self-employment income, and household income. In the final section, I offer conclusions and suggest areas of possible future research.

2 Measuring Income Inequality

I use several measures of inequality in the present paper but I perform the Bootstrap analysis only for three of them. The first is the well-known Gini coefficient, which main advantages compared with other measures of inequality are its widespread use in empirical work and its easy interpretation. Additionally, the Gini coefficient satisfies all the basic conditions of acceptable inequality indicators, including the axioms of anonymity, scale independence, population independence, and the transfer principle (Fields, 2001). Several formulas are used to calculate the Gini coefficient, but I use the one found in Athanosopulos and Vahid (2003), in which sampling weights are used in order to compensate for the fact that the sample was collected by a non-random method:

$$Gini = 1 + \frac{\sum_{j=1}^{n} y_j p_j^2}{\sum_{j=1}^{n} y_j p_j} - \frac{2}{\sum_{j=1}^{n} y_j p_j} * \sum_{j=1}^{n} \left(y_j p_j * \sum_{i=1}^{j} p_i \right)$$

$$\text{where} \quad y_1 \geq y_2 \geq .... \geq y_n. \tag{1}$$

In this expression, y is the income variable, the p_j:s are the normalized weights such that $p_j = w_j / \sum_{i=1}^{n} w_i$, and n is the sample size.

Another advantage is that the Gini coefficient allows a decomposition of the household income inequality by income source. If the household income is composed of K different components, the household income can be written as $Y = \sum_{k=1}^{K} Y_k$ and the Gini coefficient can be defined as:

$$Gini = \sum_{k=1}^{K} S_k * R_k * Gini_k. \tag{2}$$

The different components of this expression are defined as follows:

$$\begin{aligned} S_k &= \mu_k / \mu, \\ R_k &= \frac{Cov(Y_k, \rho(Y))}{Cov(Y_k, \rho(Y_k))}, \\ \rho(Y) &= \text{rank ordering of total household income,} \\ \rho(Y_k) &= \text{rank ordering of income source } k, \\ Gini_k &= \text{Gini coefficient for income source } k \text{ when all} \\ &\quad\ \text{households are taken into account.} \end{aligned}$$

Equation (2) indicates that the contribution of the inequality of income source k is made of three different components: the share of component k on total income, captured by $S_k = \mu_k / \mu$; the inequality of income source k (taking into account all households even them not receiving that income source), $Gini_k$; and a measure of the correlation of income source k and total income, R_k.[2] While S_k and $Gini_k$ only attain values in the interval $[0, 1]$, R_k can attain any value between -1 and 1. When R_k is less than zero, income source k is negatively correlated with total household income implying that this income source has a decreasing effect on total income inequality. The contribution of source k is denoted as C_k, therefore the relative contribution of the income source to the total inequality of household income is defined as

$$\Pi_k = \frac{C_k}{Gini}. \tag{3}$$

A disadvantage of the Gini coefficient, however, is that it is not possible to be decomposed in such a way that its value is exactly equal to the sum of the inequality within and between a set of mutually exclusive sub-groups of individuals. This decomposition is very informative when one is interested in explaining the level and change of inequality. An alternative measure of inequality that makes this type of decomposition possible is the Theil inequality index. This inequality indicator is a special case of the Generalized entropy family of inequality indices that when restricting its parameter to one its expression becomes

$$Theil = \sum_{j=1}^{n} p_j * \left(\frac{y_j}{\mu}\right) * \ln\left(\frac{y_j}{\mu}\right), \tag{4}$$

where μ is overall mean income, and p_j is defined as above. Besides satisfying the basic condition of acceptable inequality indicators, the Theil index possesses the property of additive decomposability, which implies that the expression above can

[2] In $\rho(Y)$ individuals are ordered in ascending order according to their household income per adult equivalent. In $\rho(Y_k)$ individuals are ordered in ascending order according to household income source k per adult equivalent.

be decomposed into a within-group ($Theil-W$) and a between-group ($Theil-B$) component, using N mutually exclusive sub-groups.

$$
\begin{aligned}
Theil &= (Theil-W)+(Theil-B), \\
(Theil-W) &= \sum_{i=1}^{N} P_i * \frac{\mu_i}{\mu} * Theil_i, && (5)\\
(Theil-B) &= \sum_{i=1}^{N} P_i * \frac{\mu_i}{\mu} * \ln\frac{\mu_i}{\mu}. && (6)
\end{aligned}
$$

P_i represents the population proportion of group i; $Theil_i$ is inequality within sub-group i; and μ_i/μ is sub-group i relative mean income. Equation (5) represents the within-group inequality, while equation (6) is the part accounted for between-groups inequality reflecting the inequality contribution that arises because of differences on the mean across sub-groups.

The third measure of inequality for which I perform a Bootstrap analyzes is the variance of log-earnings

$$Log-variance = \sum_{j=1}^{n} p_j * (\ln y_j - \overline{\ln y})^2. \qquad (7)$$

Finally, I also report the the Entropy measures with parameter 0, E(0), with parameter 2, E(2), and, in order to get a broad picture of Chile's inequality, the ratio of the ninth and first decile, P90/P10. The reader interested in the Bootstrap method used in the paper is referred to the detailed explanation given in Appendix 1 of this paper.

3 Data and Income Variables

The data used in this paper consist of seven rounds of CASEN, a nationally representative cross-sectional household survey conducted every two years from 1990 to 2000, and again in 2003 by Chile's Ministry of Planning (MIDEPLAN). The sampling scheme is not random; instead, a multistage random methodology is used to select a number of households representative of Chile as a whole. To correct for the non-random nature of the sample, each observation is assigned a number, or "weight"; these weights are used in all calculations reported in this paper. The total number of sampled households more than doubled during the research period, from 25,793 in 1990 to 71,321 in 2003. The total number of individuals surveyed increased from 105,189 to 271,716 over the same timeframe. The number of surveyed live-in domestic servants, who are excluded from this analysis due to insufficient information concerning their households, fluctuated between 376 and 630.

MIDEPLAN cooperates with Economic Commission for Latin America and the Caribbean (ECLAC) to correct the collected data. Specifically, ECLAC corrects for missing labor income, pension income, and imputed rent. Wage and salary workers showed the lowest rate of missing data (2.76%-5.10%), while employers constituted the group with the highest rate of unavailable data (3.71%-6.38%), (ECLAC, 1995). Missing labor incomes and missing pension, retirement pension, and widow pension were corrected by mean value imputation. In the correction for missing imputed rent, the missing value was replaced by a randomly selected value from the group with similar characteristics (Hot deck method). Unfortunately, ECLAC provide little detailed information about the decision to use the imputation methods mentioned previously given that other methods, such as the regression method and the conditional distribution method, are also available. Moreover, there is no scientific analysis of the affects of different correction methods on inequality indicators in Chile.

ECLAC corrected for underreporting comparing aggregate income variables from CASEN with their equivalent in the national accounts, of which the latter was assumed be the correct source of information. The ratio between the aggregated value in the two sources was the factor of correction used to correct CASEN. The aggregated value of wage and salaries from CASEN was very close to its counterpart in the national accounts, but in 1990 the discrepancy was as much as 20%. Problems of underreporting are more prevalent for self-employment incomes. In fact, in most years, self-employed workers had a factor of correction of about 1.9, indicating that the self-employment income in the national accounts was almost twice that reported in CASEN.

When I analyze household income, I use the monetary income of the household divided by an adult equivalent scale developed by Contreras (1996). The reason for using monetary income is the fact that total income may be subject to measurement errors associated with the calculation of imputed rent.[3] In my analysis, the occupations of the surveyed individuals define the variables of wages and salaries and independent labor income. Wages and salaries are defined as the monthly earnings from the principal occupation of blue-collar workers, white-collar workers, and domestic servants. Self-employment income is the monthly earnings of employers and own-account workers from their principal occupation. Due to their lack of labor income, I exclude unpaid family workers from the preceding two income definitions.

[3]For year 2000, on average, total income was 11% higher than monetary income.

4 The Chilean Economy during the 1990s

The democratic elected government that came to power in March of 1990 inherited an economy with an uninterrupted five-year period of high growth rate, a high rate of inflation, a low level of investment,[4] and a poverty rate and level of income inequality that were distressingly high (Figure 1).

The first half of the 1990s was characterized by continued high rates of GDP-growth, most of the years well above the average growth of the period, 5%, (Figure 2). The most dynamic sectors were those in which Chile has comparative advantages, such as fishing, fruit farming, and wood and paper products. Other expansive sectors included transport and communication; commerce, restaurants, and hotels; financial services; and construction, all of which had annual growth rates of more than 8% (Álvarez and Fuentes, 2004). Consequently, employment expanded by more than 10% between 1990 and 1993, while unemployment declined to nearly 7% in the 1992-93 period. Another result of the rapid economic expansion was an increase in tax collection which in turn allowed social spending to double between 1990 and 2000 of which outlays on the five most important monetary subsidies, such as assistance pensions, subsidies to poor families, and unemployment benefits, increased by 60%.

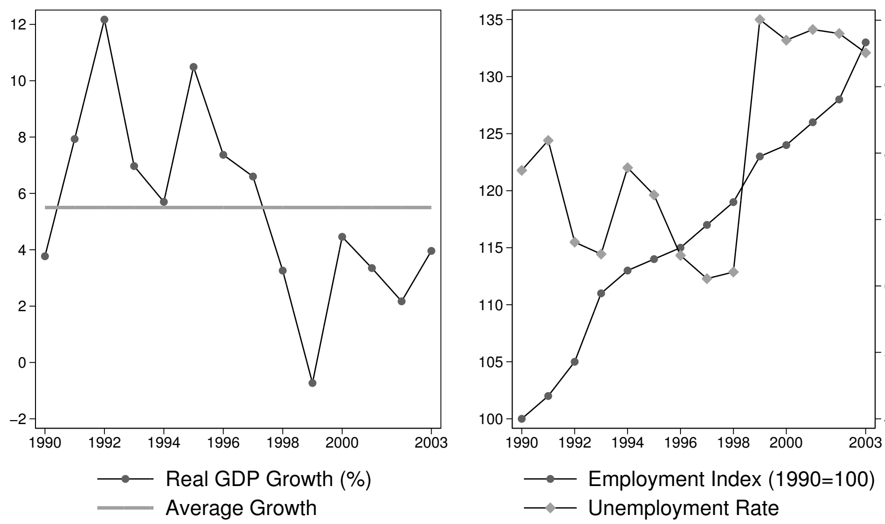
After 1995 the Chilean Peso appreciated in real terms and imbalances in the current account emerged, at the same time GDP-growth started to decline. Despite that, unemployment declined to extraordinary low levels in the whole 1996-1998 period. But, as GDP-growth rate continued to decline, and ultimately turned negative in 1999, unemployment escalated to near 10%. This collapse was caused by a private-sector spending boom; a generous public sector wage adjustment; a sharp increase in the minimum wage; a foreign-financed lending boom; and an expansionary monetary

[4] 30% and 19%, respectively, on average for the period from 1985 to 1989.

policy that created strong demand pressure in the non-tradable market, a real exchange appreciation of the Peso, and a loss of international competitiveness (Corvo and Tessada, 2003). The deceleration of output was particularly detrimental to construction; commerce, restaurants, and hotels; manufacturing; and public utilities. However, primary sectors and transport and communication continued to grow at rates near 5%.

Figure 2

Some Economic Indicators for the Chilean Economy, 1990-2003



Source: Banco Central de Chile.

In addition to the macroeconomic events already described, the research period was marked by several changes in Chile's socioeconomic composition. While the percentage of household heads older than 35 years increased from over 76% to 83%, the proportion of female household heads increased from 20% to 25%, columns (1)-(3) in Table 1. The population distribution by region is characterized by a high concentration (70%) of household heads to the four most populated regions (Metropolitan, V, VIII, and X), and by a remarkably stability over the years. Additionally, while the share of self-employed household heads declined slightly over the years, the percentage of household heads working as wage and salary workers increased by four percentage points in the first half of the research period before it declined to 50% in year 2003.

Also among wage and salary workers, the age structure shifted towards individuals older than 35, from 43% in 1990 to 54% in 2003, columns (4)-(5) in Table 1. Female labor force participation grew to almost 38% of the total wage and salary worker population over the research period. Also during the 1990s, there was a 4 percentage point increase in part-time employment, most likely as the result of the increased participation of women in the labor force. The number of white-collar workers, such as teachers, doctors, accountants, sellers, and shop assistants, increased by 8 percentage points during the period of analysis. At the same time there was a 10% points increase in the population share of of employees in service sectors and a concomitant decline in the share employed in primary and industry sectors.

The socioeconomic composition of self-employed workers differed in several respects from that of wage and salary workers, columns (7)-(9) in Table 1. The self-employed had a higher percentage of older, male, and province-resident. Moreover, the group of individuals in this group working fewer than 30 hours per week was considerably larger, and the share of white-collar occupations was lower than that of wage and salary workers. Besides the 6% increase in the proportion of part-time work, the shift towards white-collar occupations and service sectors was less pronounced than it was for wage and salary workers.

Perhaps one of the most marked changes during the research period is found for education. As a result of the growing level of education, fewer household heads were low-educated; a group that decreased from over 52% in 1990 to less than 42%

Table 1

Socioeconomic Structure by Group, 1990-2003
(All numbers represent %)

Variable	Household Heads			Wage and Salary Workers			Self-employed Workers		
	1990	1996	2003	1990	1996	2003	1990	1996	2003
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Age									
15-24	3.79	2.89	2.38	20.64	17.75	14.74	9.74	6.73	6.08
25-34	20.16	19.03	14.99	34.86	32.75	29.13	23.29	19.78	17.06
35-44	22.64	25.40	25.40	22.85	27.03	28.19	24.32	26.81	27.22
45-54	20.26	20.22	22.89	14.40	14.48	18.51	21.18	24.50	25.63
55-64	16.48	14.52	16.00	6.01	6.34	7.77	14.83	13.88	15.72
≥ 65	16.67	17.93	18.34	1.23	1.66	1.66	6.64	8.30	8.29
Gender									
Male	79.80	78.03	74.06	67.14	65.53	62.13	74.38	71.86	68.94
Female	20.20	21.27	25.94	32.86	34.47	37.87	25.62	28.14	31.06
Region[a]									
V	10.96	10.44	10.50	9.71	9.77	9.96	9.22	9.22	9.57
VIII	12.61	12.67	12.22	11.51	11.37	10.90	12.41	11.11	10.42
X	6.84	6.89	7.00	6.46	6.12	5.87	7.45	7.69	8.47
Metropolitan	39.85	40.40	40.30	45.34	45.31	45.35	39.21	40.72	40.76
Education									
≤ 8	52.24	46.41	41.16	34.77	29.33	23.66	49.43	45.21	37.21
9-12	30.07	35.44	36.51	38.39	43.72	45.96	33.16	38.08	41.38
13-16	9.72	9.54	12.44	16.83	16.33	18.25	10.22	8.87	12.82
≥ 17	7.97	8.61	9.89	10.01	10.62	12.13	7.19	7.83	8.59
Participation									
Non-working	29.76	26.19	28.18						
Part-time	4.81	5.81	7.13	5.30	7.53	9.88	22.74	21.25	28.17
Full-time	65.43	68.00	64.69	94.70	92.47	90.12	77.26	78.85	71.83
Employment Status									
Non-working	29.76	26.19	28.18						
Self-employed	22.30	21.52	21.15						
Wage and Salary	47.94	52.29	50.67						
Occupation[b]									
Blue-collar				59.83	54.58	51.48	60.16	58.97	56.67
White-collar				40.17	45.42	48.52	39.84	41.03	43.33
Sector									
Primary				18.35	15.94	14.32	19.39	19.82	15.88
Industry				28.51	26.36	22.48	22.93	20.48	24.21
Service				53.14	57.70	63.20	57.68	59.70	59.91

Source: Author's calculations from CASEN.

Notes: [a] I report only four of Chile's 13 regions, [b] For occupation and sector the values represent 1992 instead of 1990 since the classification of occupation and sector for 1990 differs from that of the other years.

in 2003. On the other hand, the group of household heads with post secondary education, that is with 13-16 or $\geq$ 17 years of formal schooling, increased from 17% to over 22% over the same period. Among wage and salary workers and the self-employed this group increased to over 30% and 20%, respectively.

Education in Chile went through a major reform in 1980 when an increased provision of the private sector was permitted at the same time that public expending on this sector declined throughout the decade. As a result, at all educational levels private or semi-private alternatives started to compete with their public counterparts. At the same time, tuition fees were introduced in higher education which permitted an expansion in places available. In addition to this, a new program of mean-tested loans for students in public universities become available to cover part or all of the tuition costs. The results was a sharp increase in the net enrollment ratio for secondary education, from 65% in 1980 to 78% in 1990. For tertiary education from 9% to 20%, respectively.

The return to democracy implied an increased flow of found to education at all levels, from 2.4% of GDP in 1990 to over 4% in 2001, and several reforms aimed to increase the quality of education and status of teachers. Net enrollment ratios for secondary and tertiary education went up to 87% and 28%, respectively. During the 1990s there was also an increased awareness that a result of the reforms introduced in the 1980s was a highly social-stratified educational system with a overwhelming proportion of students in municipal schools coming from low-income households, 80% for primary schools and 70% for secondary schools. Moreover, student loans are not provided for student in private universities, professional institutions, and technical training centers although these students are to a large extent drawn from lower socio-economic groups.

5 Wage and Salary Inequality

Despite the changing population structure of wage and salary workers and the rapid economic growth of the early 1990s, inequality in wages and salaries was quite stable over the period from 1990-96 according to the Gini coefficient. Other standard inequality measures, Theil index, E(0), and E(2) don't report such stability, instead they found 1994 to have a lower inequality than 1990. On the other hand, all inequality indicators agree in that inequality in 2000 was higher than in 1994. 4.6% higher according to the Gini coefficient, 13.7% higher according to the Theil index, and 13.0% according to the Log-variance. Moreover, while there is no clear evidence that inequality was lower in 1996 than in 1990 since the result depends on the indicator used, all indicators found wages and salaries in 2003 to be more unequally distributed than in 1990 and 1996.

Table 2

Inequality Indicators for Wages and Salaries, 1990-2003

Inequality Measure	1990 (1)	1992 (2)	1994 (3)	1996 (4)	1998 (5)	2000 (6)	2003 (7)
Gini	0.455	0.458	0.451	0.454	0.462	0.472	0.458
SE	(0.006)	(0.004)	(0.004)	(0.006)	(0.006)	(0.007)	(0.006)
Theil	0.437	0.439	0.407	0.416	0.460	0.463	0.452
SE	(0.017)	(0.010)	(0.009)	(0.016)	(0.037)	(0.025)	(0.039)
Log-variance	0.619	0.589	0.592	0.625	0.611	0.669	0.632
SE	(0.011)	(0.008)	(0.008)	(0.010)	(0.010)	(0.018)	(0.008)
E(0)	0.359	0.357	0.345	0.354	0.365	0.387	0.366
E(2)	1.940	1.801	1.493	1.656	3.546	2.028	3.920
P90/P10	5.54	6.00	6.25	6.00	6.67	6.43	6.40

Source: Author's calculations from CASEN.

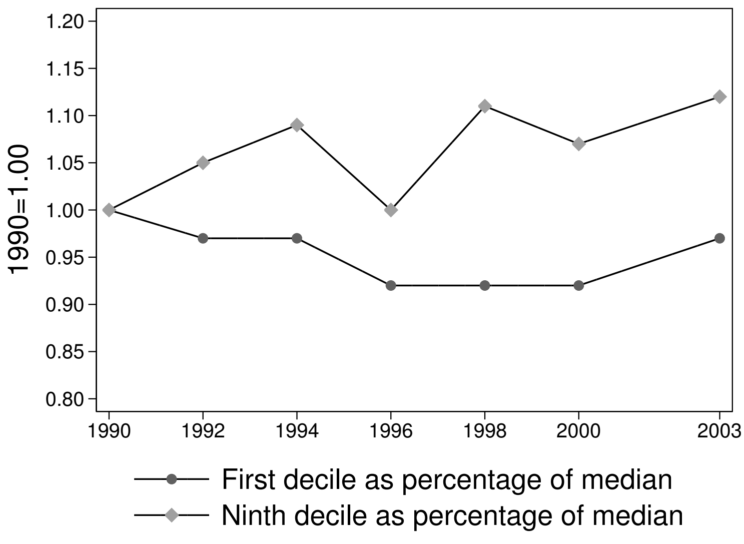
Notes: Standard errors, *SE*, are calculated using 500 Bootstrap replications.

The Bootstrap analysis of these inequality changes (Tables A1-A3 in the appendix), indicates that the Gini coefficient, the Theil index, and Log-variance found 2000 to have a significantly higher level of inequality than 1994. No other change was found to be significant by all inequality measures analyzed. Therefore, we can only conclude that there was an inequality deterioration between 1994 and 2000.

Additional information is obtained looking at the inequality in the different parts of the distribution of wages and salaries. Figure 3 presents the ninth decile and first decile as percentage of the median. To make the two series comparable I set 1990=1.0. According to this figure, inequality in the lower part of the distribution in the period 1996-2000 was higher than in other years.

Figure 3

Inequality of Wages and Salaries in Different Parts of the Distribution, 1990-2003



Source: Author's calculations from CASEN.

In the upper part of the distribution, inequality increased rapidly in 1992 and 1994. But after a momentary decline in 1996, the inequality in this part of the distribution returned to its 1994-level. In consequence, Chile ended the decade with an inequality that was clearly worse than in 1990, both in the upper and lower part of the distribution of wages and salaries.

The decomposition of the Theil index by population sub-group indicates that the between-group component of wage and salary inequality was important in only education and occupation, explaining over 40% (0.190) and 18% (0.085), respectively in 2000 (Table A4 in the appendix). The underlying reason for the pronounced role of these variables is the great difference between the relative income of the best and worst paid groups. For instance, in 1996 the ratio between the relative income of these two groups amounted to 5.0 for education and 2.2 for occupation, while it was below 2 for most other groups.

While the within-education component remained relatively stable over the years, the between-education component generally increased, almost doubling between 1990 and 2000; thus, the between-education component's explanatory power increased from 27% in 1990 to 40% in 2000. The picture for occupation does not mirror that of education. Occupation had a relatively stable between-group component over the whole period. The single exception to that stability came in 1994, when it declined from 0.076 to 0.062 before increasing to 0.080 in the following years. The contributions of age and region were also stable, but, at 5%, were smaller in relative size. At the same time, gender, participation, and sector contributed only negligibly.

Despite that the total inequality change of wages and salaries in the periods 1996-90, 2003-1996, and 2003-1990 were no significant, education had significant, relatively large, and positive between-group components in the inequality changes in both sub-periods which more than compensated the decline in the inequality within groups in the 1990-96 period. The between-group component of participation was also significant, but very small. That sector had almost no contribution to the level of and change in inequality is rather surprising considering the population share of service sectors increased by nearly 10 percentage points during the research period; however, given that differences in relative wages and salaries across sectors are much smaller than those across educational levels or occupations, this result makes more sense.

Table 3

Changes in Theil Index ($\Delta Theil$) of Wages and Salaries, 1990-2003

(Theil-W and Theil-B represent within and between-group component)

	Age	Gender	Region	Education	Participation	Occupation	Sector
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				1996-1990			
$\Delta Theil - W$	-0.024	-0.024	-0.034	-0.049*	-0.028	*n.a.*	*n.a.*
(95% CI)	(-0.065;0.018)	(-0.065;0.021)	(-0.075;0.010)	(-0.081;-0.012)	(-0.071;0.018)	*n.a.*	*n.a.*
$\Delta Theil - B$	-0.002	-0.002	0.008*	0.023*	0.002	*n.a.*	*n.a.*
(95% CI)	(-0.007;0.005)	(-0.006;0.001)	(0.003;0.014)	(0.005;0.044)	(-0.001;0.005)	*n.a.*	*n.a.*
$\Delta Total$	-0.026	-0.026	-0.026	-0.026	-0.026	*n.a.*	*n.a.*
				2003-1996			
$\Delta Theil - W$	0.036	0.037	0.042	0.013	0.030	0.040	0.038
(95% CI)	(-0.053;0.098)	(-0.059;0.099)	(-0.052;0.102)	(-0.077;0.074)	(-0.069;0.094)	(-0.049;0.098)	(-0.057;0.101)
$\Delta Theil - B$	0.002	0.001	-0.004	0.025*	0.007*	-0.002	0.000
(95% CI)	(-0.005;0.005)	(-0.003;0.004)	(-0.011;0.003)	(0.009;0.040)	(0.003;0.012)	(-0.012;0.007)	(-0.004;0.003)
$\Delta Total$	0.038	0.037	0.038	0.038	0.037	0.038	0.038
				2003-1990			
$\Delta Theil - W$	0.013	0.014	0.008	-0.036	0.003	*n.a.*	*n.a.*
(95% CI)	(-0.076;0.076)	(-0.080;0.082)	(-0.082;0.075)	(-0.129;0.024)	(-0.091;0.072)	*n.a.*	*n.a.*
$\Delta Theil - B$	0.000	-0.001	0.004	0.048*	0.009*	*n.a.*	*n.a.*
(95% CI)	(-0.008;0.007)	(-0.005;0.002)	(-0.002;0.010)	(0.031;0.067)	(0.006;0.013)	*n.a.*	*n.a.*
$\Delta Total$	0.013	0.013	0.012	0.012	0.012	*n.a.*	*n.a.*

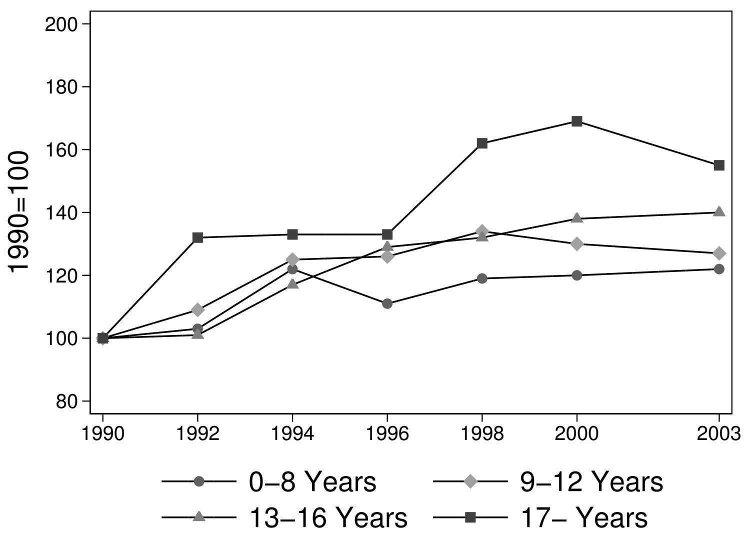
Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications.* Represents that the inequality change was significant. *n.a.* the decomposition cannot be done for these years since the classification of occupation and sector of 1990 is different from that of the other years.

A closer inspection of the average wage and salary at different educational levels reveals that the strong economic expansion of 1990-92 largely benefited the most educated individuals, for whom the average wage and salary increased by nearly 40% (Figure 4). However, in 1994, the year with lowest inequality, the wage and salary of the most educated stagnated, while the less well-educated experienced a rapid pay growth before a decline in 1996. In consequence, the relative wage and salary of the most educated increased from 2.43 to 2.52 while it declined from 0.61 to 0.52 for the least educated.

Figure 4

Wage and Salary Trends by Years of Schooling, 1990-2003



Source: Author's calculations from CASEN.

Nevertheless, after 1996, the incomes of those with the greatest level of education experienced another strong expansion; by the end of the decade, the average wage and salary for this group was 70% higher than at the beginning. In contrast, those with 9-12 years of schooling had an almost constant average wage and salary after 1994 and therefore a declining relative average income, from 0.82 in 1996 to 0.72 in year 2003. In conjunction with the increase in the population share of this group, from 38% in 1990 to 45% in 2003, the decline in their relative income diminished the between-group component and counteracted the large positive contribution of the most educated.

The heterogenous pattern of earnings across educational groups was most likely even accompanied by different growth rates within economic sectors. To study this possibility the four educational levels are interacted with the three different industries.[5] Highly educated employees in all economic sectors drove a large portion of the between-group component of the change in inequality; they did not, however, all contribute in the same manner. The highly educated employees in service sectors contributed to the increase in the between-group component through a higher population share, from 7.6% of the total of wage and salary workers in 1994 to 9.9% in 2000, since their μ_i/μ was virtually constant over the years. Highly educated employees in the primary and industry sectors, on the other hand, contributed through a strong relative wage and salary growth. This increase, totaling over 30%, solidified the status of these groups as the best paid among all wage and salary workers. This pattern among the highly educated falls in line with Chile's general economic growth in the second half of the 1990s; service sectors reported lower growth rates during this period, while primary and some industry sectors continued to grow at rates in excess of 5%.

6 Self-employment Income Inequality

The value of the Gini coefficient, Theil index, E(0), and E(2) for self-employment incomes were lower in the period 1996-1998 than in other years. While the Gini co-

[5] Results not reported here.

efficient indicate that this decline amounted to 3% of its 1990 value, the Entropy measures of inequality suggest a decline of 10%-15%. P90/P10 and Log-variance, on the other hand, indicate that year 1998, as well as 1992, had 10%-15% lower levels of inequality than 1990. Moreover, all measures of inequality report a deterioration of inequality in 2003 compared with 1996 and 1990.

Table 4

Inequality Indicators for Self-employment Incomes, 1990-2003

Inequality Measure	1990 (1)	1992 (2)	1994 (3)	1996 (4)	1998 (5)	2000 (6)	2003 (7)
Gini	0.615	0.610	0.626	0.597	0.592	0.623	0.624
SE	(0.007)	(0.008)	(0.007)	(0.010)	(0.007)	(0.016)	(0.008)
Theil	0.824	0.811	0.841	0.722	0.729	0.835	0.862
SE	(0.030)	(0.032)	(0.035)	(0.037)	(0.024)	(0.065)	(0.042)
Log-variance	1.270	1.144	1.317	1.218	1.141	1.298	1.342
SE	(0.029)	(0.025)	(0.024)	(0.033)	(0.026)	(0.044)	(0.028)
E(0)	0.717	0.686	0.746	0.667	0.647	0.737	0.752
E(2)	4.306	4.194	4.494	3.154	3.637	4.393	5.148
P90/P10	15.12	12.90	17.50	13.33	12.88	18.25	16.67

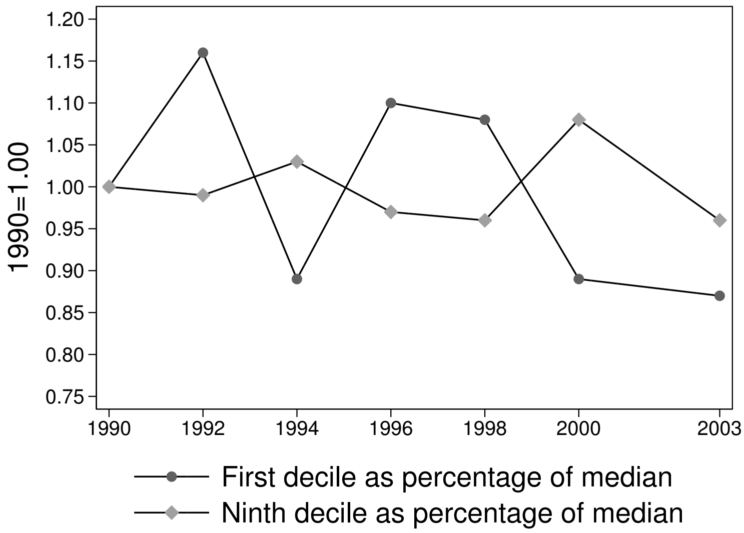
Source: Author's calculations from CASEN.
Notes: Standard errors, *SE*, are calculated using 500 Bootstrap replications.

The Bootstrap analysis of inequality changes (Table A5-A7 in the appendix) suggest a significant distributional improvement in 1998, compared with 1990 and 2003, according to the Gini coefficient and Log-variance. The Theil index, on the other hand, indicates that even in 1996 inequality was significantly lower than in these two years. Moreover, while only the Theil index found 1996 to have a significantly lower inequality than 1990, all inequality indicators found a significant change between 1996 and 2003. In consequence, there is evidence of an inequality decline in 1998 with respect to 1990 and of a worse distribution of self-employment incomes in the period 2000-2003 compared with 1996-98.

The explanation to the lower inequality in 1996-1998 and 1992 is found in Figure 5. According to this figure, in 1992 and 1996-1998 the inequality in the lower part of the distribution was lower than in other years at the same time as the inequality in the upper part was relatively stable. This picture totaly changed in year 2000 when inequality in the lower part of the distribution declined sharply at the same time as it increased in the upper part.

Figure 5

Inequality of Self-employment Incomes in Different Parts of the Distribution, 1990-2003



Source: Author's calculations from CASEN.

This resulted in a level of inequality that was higher than in most other years. According to P90/P10, inequality declined by 15% in year 1992 and by 25% in 1996. The increase in 2000 was as much as 40% compared with 1998. The lower level of inequality in the lower part of the distribution in 1992 and in the 1996-1998 period coincide quite well with low rates of unemployment in Chile, see Figure 2. This might be an indication that when the labor market conditions are good, the less productive self-employed workers leave this segment of the labor market and thereby induce an inequality decline in this part of the distribution.

As it was the case in the previous section, the most important variable for the decomposition by population sub-group is education, for which the between-group component explains over 19% of the Theil index (Table A8 in the appendix). But while the ratio of the relative income of the group with highest and lowest level of education was 5.0 for wages and salaries, for self-employed workers it was nearly 7.0. For participation and region ($Theil - B$) explains roughly 4% of the index. However, although similar in trend to the last section, the self-employment results differ in scale. For wages and salaries in 2000, the contributions of education and participation were 0.190 and 0.013, whereas for self-employment incomes their contributions were 0.306 and 0.051, respectively. Moreover, these variables had a clear tendency to an increasing between-group component. In fact, the share explained by the between-group component of education increased from 19% in 1990 to 37% in 2000.

The between-group component of age accounts for less than 5% of the total, while income differences between genders and sectors had no explanatory power. The classification for the variables of occupation and sector was different in 1990, and therefore no decomposition could be completed for this year. On the other hand, for 1996 and 2003, income differences between groups explained 20% and 3%, respectively, implying that occupation was the second most important variable in explaining the level of inequality of self-employment incomes.

In the face of the large overall decline in the dispersion of self-employment incomes between 1990 and 1996, only education and region showed a large and significant between-group inequality increase. In the following period, the between-group component of education was again positive and statistically significant, accounting for near 70% of the change. Participation, accounting for 20% of the change, was also significant. For the entire period 1990-2003 the only variables for which the between-group change was significant were region, education, and participation which all had an inequality increasing effect and thereby counteracted the decline of inequality within groups that emerged between 1990 and 1996.

Table 5

Changes in Theil Index ($\Delta Theil$) of Self-employment Incomes, 1990-2003

(Theil-W and Theil-B represent within and between-group component)

	Age (1)	Gender (2)	Region (3)	Education (4)	Participation (5)	Occupation (6)	Sector (7)
				1996-1990			
$\Delta Theil - W$	-0.105*	-0.096	-0.140*	-0.170*	-0.113*	*n.a.*	*n.a.*
(95% CI)	(-0.211;-0.006)	(-0.206;-0.001)	(-0.246;-0.042)	(-0.279;-0.073)	(-0.219;-0.018)	*n.a.*	*n.a.*
$\Delta Theil - B$	0.000	-0.008*	0.036*	0.067*	0.009	*n.a.*	*n.a.*
(95% CI)	(-0.015;0.0015)	(-0.019;0.004)	(0.014;0.056)	(0.018;0.114)	(-0.007;0.026)	*n.a.*	*n.a.*
$\Delta Total$	-0.105	-0.104	-0.104	-0.103	-0.104	*n.a.*	*n.a.*
				2003-1996			
$\Delta Theil - W$	0.134*	0.124	0.118*	0.042	0.110*	0.126*	0.144*
(95% CI)	(0.029;0.234)	(0.008;0.225)	(0.016;0.212)	(-0.052;0.147)	(0.004;0.213)	(0.031;0.218)	(0.033;0.251)
$\Delta Theil - B$	0.003	0.013*	0.019	0.095*	0.026*	0.011	-0.007
(95% CI)	(-0.014;0.019)	(0.001;0.025)	(-0.004;0.044)	(0.040;0.146)	(0.009;0.043)	(-0.019;0.042)	(-0.018;0.004)
$\Delta Total$	0.137	0.137	0.137	0.137	0.136	0.137	0.137
				2003-1990			
$\Delta Theil - W$	0.029	0.028	-0.022	-0.128*	-0.003	*n.a.*	*n.a.*
(95% CI)	(-0.008;0.128)	(-0.083;0.127)	(-0.134;0.080)	(-0.218;-0.034)	(-0.114;0.101)	*n.a.*	*n.a.*
$\Delta Theil - B$	0.004	0.005	0.055*	0.161*	0.035*	*n.a.*	*n.a.*
(95% CI)	(-0.011;0.018)	(-0.005;0.017)	(0.035;0.074)	(0.105;0.220)	(0.018;0.054)	*n.a.*	*n.a.*
$\Delta Total$	0.033	0.033	0.033	0.033	0.032	*n.a.*	*n.a.*

Source: Author's calculations from CASEN.

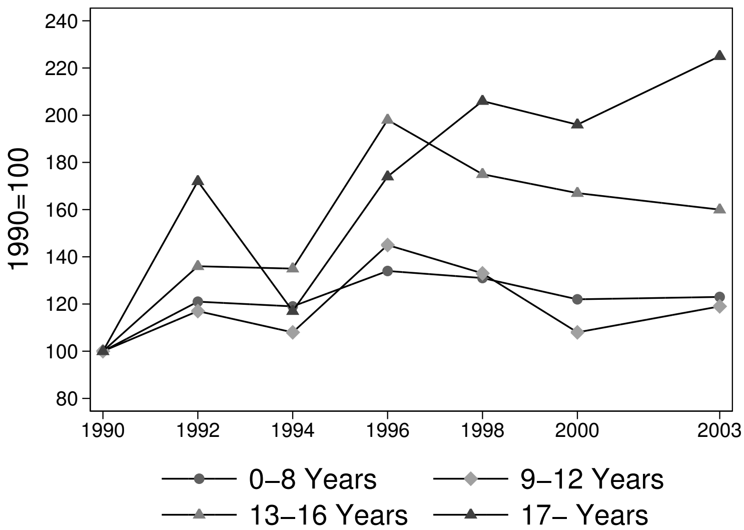
Notes: Calculated using 500 Bootstrap replications.* Represents that the inequality change was significant. *n.a.* the decomposition cannot be done for these years since the classification of occupation and sector of 1990 is different from that of the other years.

Since the second half of the research period is characterized by a proliferation of part-time employment, and the income differences between these groups increased over time, the increased explanatory power and significance of participation is partially expected. Part-time employment was practically unchanged for self-employed workers in the first half of the research period despite the continuous increase in female labor participation; thus, the increased role of part-time employment was most likely an effect of the weakened labor market at the end of the decade.

Further study of this income variable reveals that, when comparing 1996 to 1990, the self-employment incomes of those with 13-16 or more than 17 years of schooling increased more than the incomes of those with less education; this explains the large contribution of the between-education inequality in this period. In the years following, only the most educated achieved an additional increase in average income. The income of other groups was stable or even declined. The overall result was a between-education inequality that explained almost the totality of the inequality change during the research period.

Figure 6

Self-employment Income Trends by Years of Schooling, 1990-2003



Source: Author's calculations from CASEN.

To shed light on those sectors contributing the most to the between-group component of education, I created a more detailed classification of self-employed workers.[6] Education's large contribution to the inequality change of the first sub-period was driven by service-sector workers with 13-16 and more than 17 years of education. In the following sub-period, in addition to well-educated service-sector workers, the most-educated workers in industry sectors contributed the most. For both groups, relative self-employment income, rather than change in population share, explained the high contribution to inequality change.

7 Household Income Inequality

The value of all household income inequality indicators declined in 1994, compared with 1990, and deteriorated in 2000 compared with 1994; 5.8% higher according to the Gini coefficient, 16.7% higher according to the Theil index, and 11.5% according to the Log-variance. On the other hand, no conclusion could be drawn comparing 1996 with 1990, 2003 with 1996, and 2003 with 1990 since not all indicators moved in the same direction.

The Bootstrap analysis of these inequality changes (Tables A10-A12 in the appendix) indicates that, since the 95% interval of the change in the Gini coefficient, the Theil index, and the Log-variance between 1994 and most of its following years do not include zero, the inequality changes between these years are significant. Nevertheless, the change in inequality between 1990 and 1994 was significant only according to the Gini coefficient and Log-variance. Putting these results together, there is evidence of a reduction of inequality in 1994 and a clear deterioration in inequality at least for the years 1998-2000.

[6] Results not reported here.

Table 6

Inequality Indicators for Household Income per Adult Equivalent, 1990-2003

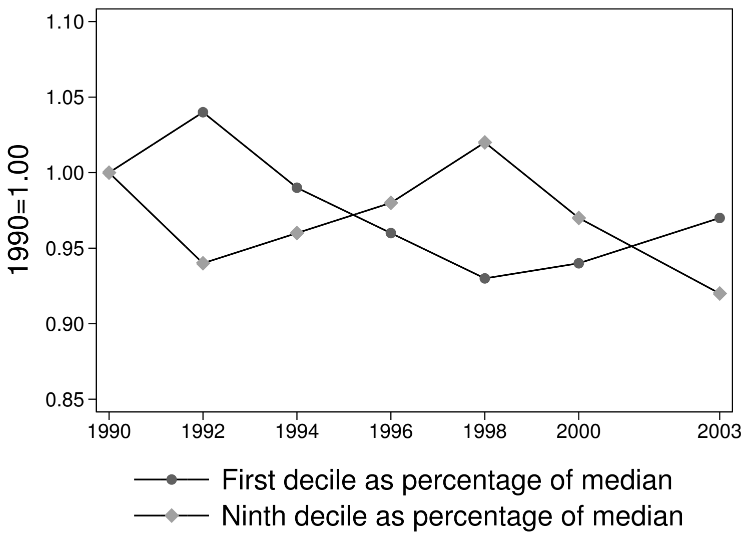
Inequality Measure	1990 (1)	1992 (2)	1994 (3)	1996 (4)	1998 (5)	2000 (6)	2003 (7)
Gini	0.546	0.539	0.530	0.548	0.554	0.561	0.545
SE	(0.002)	(0.002)	(0.002)	(0.003)	(0.002)	(0.004)	(0.003)
Theil	0.636	0.623	0.585	0.624	0.646	0.683	0.657
SE	(0.011)	(0.007)	(0.012)	(0.014)	(0.009)	(0.018)	(0.017)
Log-variance	0.923	0.844	0.880	0.954	0.995	0.982	0.904
SE	(0.007)	(0.004)	(0.005)	(0.009)	(0.011)	(0.011)	(0.005)
E(0)	0.533	0.510	0.498	0.538	0.554	0.569	0.530
E(2)	3.345	3.073	3.070	2.928	3.411	3.756	4.554
P90/P10	9.99	8.98	9.65	10.14	10.88	10.32	9.50

Source: Author's calculations from CASEN.

Notes: Standard errors, *SE*, are calculated using 500 Bootstrap replications.

Figure 7

Inequality in Different Parts of the Distribution of Household Income per Adult Equivalent, 1990-2003



Source: Author's calculations from CASEN.

A closer inspection of the inequality of household income indicates that inequality in the upper and lower part of the distribution moved in the same direction, see Figure 7. After an inequality decline in 1992, the inequality in the lower part of the distribution increased until a trend-break was reached in 1998. During these years, the ninth decile, as percentage of the median, increased by more than 8%. Therefore, the clear deterioration in the dispersion of household income between 1992 and 1998 suggested by P90/P10, which amounted to 20%, was associated with a deterioration in inequality both in the upper and lower parts of the distribution.

In contrast to previous sections, where a decomposition by population subgroup was used, this section analyzes the inequality of household income per adult equivalent by income source. Using this decomposition we are able to analyze those incomes that are dominant in the distribution of household income. In order to link the household income to the two income variables analyzed in, and the result from, previous sections, the total income of the household is divided into 11 different income components:

1. *Wage Primary Education*—Earnings from principal occupation of employees with ≤ 8 years of schooling.
2. *Wage Secondary Education*—Earnings from principal occupation of employees with 9-12 years of schooling.
3. *Wage Post-secondary Education*—Earnings from principal occupation of employees with 13-16 years of schooling.
4. *Wage University Education*—Earnings from principal occupation of employees with ≥ 17 years of schooling.
5. *Self-employment Primary Education*—Earnings from principal occupation of self-employed workers with ≤ 8 years of schooling.
6. *Self-employment Secondary Education*—Earnings from principal occupation of self-employed workers with 9-12 years of schooling.
7. *Self-employment Post-secondary Education*—Earnings from principal occupation of self-employed workers with 13-16 years of schooling.
8. *Self-employment University Education*—Earnings from principal occupa-

tion of self-employed workers with ≥ 17 years of schooling.

9. *Subsidies*—Monetary subsidies from the state to people including assistance pension, unemployment benefit, subsidy to poor families, family assignment system, and other monetary transfers.

10. *Pension*—Pension incomes.

11. *Other Incomes*—Other incomes of the households such as self-supply and consumption value of the agricultural commodities produced at home plus property revenue, interest income, allowances and bonus. It also includes labor income from other work than principal occupation.

In previous sections we found that during the 1990s the inequality between educational groups increased both among wage and salary workers and among the self-employed. The reason for this was that the labor income of the most educated increased more than for other groups. Therefore, it could be informative to analyze the effect that a marginal increase in these income sources would have in the inequality of household income. This information is given by the expression developed by Pyatt *et al.* (1980). Assuming unchanged labor and production decisions, consider an exogenous increase of income source k by a factor ϕ. Then, the derivative of the Gini coefficient with respect to a change in income source k is

$$\frac{\partial Gini}{\partial \phi} = S_k * (R_k * Gini_k - Gini), \tag{8}$$

where S_k, R_k, $Gini_k$, and $Gini$ denote factor k income share, a measure of the correlation between income source k and total income, the Gini coefficient of that income source, and the Gini coefficient of total income. In my results I report the effect on total inequality of an increase in a particular income source as a percentage of the Gini coefficient of household income, $(\partial Gini/\partial \phi)/Gini$.

Table 7

Decomposition of the Gini Coefficient of Household Income per Adult Equivalent by Income Source in 1990, 1996, and 2003

	S_k	R_k	$Gini_k$	C_k	Π_k	Percentage change in Gini coefficient
	(1)	(2)	(3)	(4)	(5)	(6)
				1990		
Wage Primary Education	0.093	0.035	0.807	0.003	0.48	-0.089
Wage Secondary Education	0.143	0.378	0.797	0.043	7.86	-0.065
Wage Post-secondary Education	0.099	0.704	0.923	0.064	11.76	0.019
Wage University Education	0.108	0.848	0.959	0.088	16.05	0.054
Self-employment Primary Education	0.083	0.482	0.922	0.037	6.73	-0.015
Self-employment Secondary Education	0.107	0.750	0.955	0.077	14.06	0.033
Self-employment Post-secondary Education	0.047	0.855	0.985	0.039	7.21	0.026
Self-employment University Education	0.063	0.929	0.989	0.058	10.58	0.044
Subsidies	0.009	-0.144	0.714	-0.001	-0.18	-0.011
Pension	0.087	0.436	0.866	0.033	5.99	-0.028
Other Incomes	0.161	0.830	0.794	0.106	19.43	0.033
Total	1.00			0.547	100	
				1996		
Wage Primary Education	0.069	-0.136	0.813	-0.008	-1.38	-0.083
Wage Secondary Education	0.161	0.336	0.748	0.040	7.36	-0.087
Wage Post-secondary Education	0.098	0.665	0.917	0.060	10.91	0.011
Wage University Education	0.118	0.844	0.953	0.095	17.30	0.057
Self-employment Primary Education	0.066	0.431	0.926	0.026	4.79	-0.019
Self-employment Secondary Education	0.126	0.760	0.943	0.090	16.45	0.041
Self-employment Post-secondary Education	0.058	0.901	0.986	0.052	9.46	0.037
Self-employment University Education	0.081	0.952	0.988	0.076	13.83	0.059
Subsidies	0.010	-0.342	0.712	-0.002	-0.42	-0.015
Pension	0.064	0.455	0.905	0.026	4.77	-0.015
Other Incomes	0.150	0.777	0.793	0.093	16.87	0.020
Total	1.00			0.549	100	
				2003		
Wage Primary Education	0.057	-0.127	0.843	-0.006	-1.12	-0.069
Wage Secondary Education	0.159	0.266	0.736	0.031	5.69	-0.102
Wage Post-secondary Education	0.109	0.659	0.910	0.066	12.04	0.011
Wage University Education	0.145	0.867	0.946	0.119	21.87	0.074
Self-employment Primary Education	0.047	0.355	0.937	0.016	2.88	-0.019
Self-employment Secondary Education	0.105	0.669	0.934	0.066	12.05	0.015
Self-employment Post-secondary Education	0.061	0.859	0.982	0.051	9.36	0.033
Self-employment University Education	0.111	0.967	0.989	0.106	19.46	0.085
Subsidies	0.011	-0.410	0.775	-0.003	-0.64	-0.017
Pension	0.055	0.426	0.913	0.021	3.94	-0.017
Other Incomes	0.140	0.703	0.804	0.079	14.48	0.006
Total	1.00			0.544	100	

Source: Author's calculations from CASEN.

The results of the decomposition are found in Table 7. One striking result is that behind the stability of the Gini coefficient of household income, 0.54 for all three years, several underlying components changed in size over the research period, implying a change in the structure of inequality. In the beginning of the research period the income source *Other Incomes* explained the largest share of the Gini coefficient, mostly thanks to its large share on total income (16%) and the high positive correlation between the rank ordering of this income source and the rank ordering of total income (0.83). The 19% explained by this income source was followed by the 16% explained by *Wage University Education* and the 14% explained by *Self-employment Secondary Education*.

Monetary subsidies had the smallest share on total income (less than 1%) and had a negligible negative contribution to total inequality (-0.18%). This negative contribution was due to the negative value of R_k, which implies that individuals receiving monetary subsidies were disproportionately found at the lower part of the household income distribution. In essence, these results indicate that boosting monetary subsidies would reduce inequality, but the reduction is only marginal. The size of the relative reduction of the Gini coefficient when monetary subsidies increases by 1% is found in the last column of Table 7. An increase of this income source by 1% reduces total inequality by only 0.011%. A much bigger inequality reducing effect is obtained increasing *Wage Primary Education*, -0.089%. In contrast, the last column of Table 7 indicates that an increase in *Wage University Education* and *Self-employed University Education* by 1% increases the Gini coefficient by nearly 0.05%.

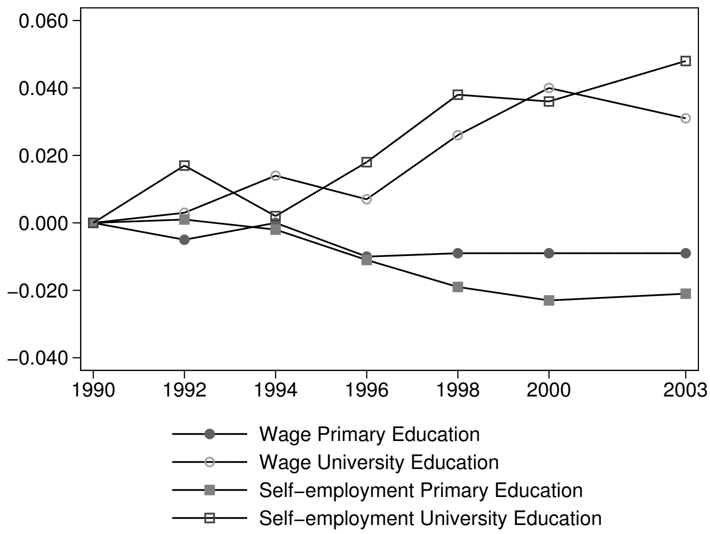
In 1996 the picture changed somewhat. First, the share of inequality explained by the income of wage and salary workers with university education rose to over 17% and became the largest contribution to explain the level of inequality of household income in that year. Since R_k and $Gini_k$ were unchanged between 1990 and 1996 for this income source, this result is explained by an increase in its income share. Other sources that increased their share were *Self-employment Secondary Education* and *Self-employment University Education*. But in contrast to the contribution of *Wage University Education*, which increased by 7%, the contribution of these income sources went up by 17% and 30%, respectively. Together, these three income sources accounted for nearly 25% of total income and 25% of the individuals had access to them, but they accounted for more than 45% of the Gini coefficient.

Second, the contribution of *Wage Primary Education* become negative, indicating that this income source had an inequality decreasing effect on total inequality. An underlying reason for this is the negative correlation between the rank of this income source and the rank of total income. That is, households with a large *Wage Primary Education* tend to be placed at the lower end of the distribution of household income per adult equivalent. It is surprising that monetary subsidies continued to have such a small contribution to explain inequality, less than 1%, given that total outlay on the five most important monetary subsidies increased by more than 60% during the 1990s. This is explained by the fact that although the share of this income source increased slightly during the research period, it continued to account for no more than 1% of total income.

In the last year of the research period the contribution of wage and salaries of employees with university education grew to over 20%. Self-employed workers with the same level of education increased to over 19%, implying that the explanatory power of this income source increased by more than 80% comparing this year with 1990. For this income source not only S_k increased over the years, even R_k went up during the research period. A high value of R_k implies that an increase of this income source would to a large degree benefit the better of an thus deteriorate the Gini coefficient. As this correlation measure grew larger over the years, the percentage change in the Gini coefficient when this income source increases by 1%, changed from 0.044% in 1990 to 0.085% in 2003. The increased importance of this income source to explain total inequality is clearly reflected by Figure 8.

Figure 8

Changes in the contribution to the Gini coefficient, 1990-2003



Source: Author's calculations from CASEN.

In this figure we can see that the change in the contribution of the different income sources explains the inequality of the years covered by CASEN, with 1990 as base year. Both *Wage University Education* and *Self-employment University Education* drew the Gini coefficient of household income upwards. This was especially important after 1996, when the income of the self-employed with university education started to increase followed by their wage and salary counterparts in 1998. On the other hand, the contribution of the wage and salary and the self-employment income of workers with 0-8 years of schooling had a negative contribution during the whole period 1996-2003. The pattern of these contributions clearly reflect the divergent pattern of the labor income of the well educated, in relation to the labor income of less educated workers, reported in Figures 4 and 6. Therefore, the large increase in the earnings of individuals with university education was not only a driving factor in the between-group component in the inequality of wages and salaries and self-employment incomes, it was also very important to explain the deterioration in the dispersion of household income found between 1994 and 2000.

Very instructive information is obtained taking advantage of the formula developed by Lerman and Yitzhaki (1994), according to whom $Gini_k = \widehat{Gini}_k * P_k + (1 - P_k)$, where $\widehat{Gini}_k$ is the inequality of source k when only the households actually receiving that particular income source are taken into account, P_k is the proportion of households with positive income source, and $(1 - P_k)$ is the population share of households without access to income source k. In the remaining of the paper I refer to $\widehat{Gini}_k * P_k$ as the source inequality component and to $(1 - P_k)$ as the source inaccessibility component.

The increased level of education in Chile is clearly reflected by columns (4)-(6) of Table 8. The percentage of individuals receiving *Wage Primary Education* and *Self-employment Primary Education* clearly declined during the research period, from 31% to 23% and from 17% to 12%, respectively. In contrast, the percentage of individuals receiving *Wage University Education* increased from 8% to 10% over the same period. But despite the reduction in its inequality, $Gini_k$ of this income source remained relatively unchanged over the years. The explanation is that the increase in the source inequality component (due to the combination of a decline in $\widehat{Gini}_k$ and an increase in P_k) was accompanied by a decline in the source inaccessibility component of the same size (due to the increase in P_k). Something different occurred with *Wage*

Primary Education, for which $\widehat{Gini}_k$ also declined over the years. In addition to this, there was a concomitant reduction in the proportion of households receiving this income source, -25%, generating a clear reduction in the source inequality component. At the same time, the increase in the source inaccessibility component was larger in size. In consequence, the dispersion of this income source among all households increased from 0.807 in 1990 to 0.843 year 2006.

Table 8

Inequality of the different Income Sources in 1990, 1996, and 2006

	$\widehat{Gini}_k$			P_k		
	1990	1996	2003	1990	1996	2003
	(1)	(2)	(3)	(4)	(5)	(6)
Wage Primary Education	0.391	0.354	0.340	0.318	0.290	0.238
Wage Secondary Education	0.414	0.402	0.383	0.347	0.420	0.427
Wage Post-secondary Education	0.477	0.462	0.475	0.148	0.155	0.171
Wage University Education	0.533	0.511	0.478	0.088	0.095	0.104
Self-employment Primary Education	0.543	0.487	0.469	0.170	0.144	0.119
Self-employment Secondary Education	0.621	0.560	0.546	0.118	0.130	0.145
Self-employment Post-secondary Education	0.571	0.562	0.574	0.036	0.031	0.043
Self-employment University Education	0.579	0.531	0.580	0.026	0.026	0.027
Subsidies	0.433	0.483	0.549	0.504	0.557	0.498
Pension	0.505	0.473	0.485	0.270	0.180	0.170
Other	0.635	0.697	0.705	0.563	0.681	0.665

Source: Author's calculations from CASEN.

Self-employment University Education, which was one of the income sources with the largest relative increase in its share on total inequality, was highly concentrated to less than 3% of the population over the whole research period. Together with this, the unchanged $\widehat{Gini}_k$, made the total inequality of this income source highly stable over the research period.

8 Conclusions

This paper analyzed the inequality of Chilean wages and salaries, self-employment incomes, and household income during the 1990-2003 period. The work complements earlier studies analyzing several issues that were previously absent from the relevant body of literature.

The Bootstrap analysis of the inequality changes suggests that for wages and salaries, there was a significant increase in inequality between 1994 and 2000. For self-employment incomes, significant changes occurred comparing 1996-98 with 1990 and comparing 2000-03 with 1996-98. Moreover, I found clear evidence of a significant deterioration in the dispersion of household income in 2000 compared with 1994.

The results indicate that, because their socioeconomic composition and inequality are quite different from those of wage and salary workers, a separate analysis of self-employed workers is justifiable. The group of self-employed workers has a higher percentage of older, male, province-resident, part-time, and low-educated workers than does the group of wage and salary workers. But for both, wages and salaries and self-employment incomes, in line with previous studies in Chile, education was found to be the single most important variable to explain inequality. The share explained by education, 20%-40%, was followed by share explained by occupation, 15%-20%, and region, 1.6%-8.0%.

Despite the changes in inequality of wage and salary in the periods 1996-90, 2003-1996, and 2003-1990 were no significant according to the Bootstrap analysis, the changes in the between-education component were positive and significant for these years, more than compensating the inequality decline in the first half part of the research period. Behind this result I found a strong increase in the average wage and salary of highly educated employees, especially in primary sectors and industry sectors. For self-employment incomes, the between-education component was large and positive (inequality increasing) over the whole research period but 50% higher in the 2003-1996 period. The strong increment in the relative self-employment income of the most educated service-sector workers drove this result.

From the decomposition of household income emerge three important results. First, behind the highly stable Gini coefficient of household income, the earnings of employees and of self-employed with university education become the two most important income sources to explain the total inequality at the same time as the earnings of their counterparts with primary education decreased their role. Second, the most important factor to explain this result is the increase in the income share of *Wage University Education* and *Self-employment University Education* over the research period. The share of these two income sources as proportion of total household income increased from 10% and 6%, respectively, in 1990 to 14% and 11% in the last year of the research period. Third, behind the apparently stable inequality of wages and salaries of employees with university education, $Gini_k$, I found a clear decline in the inequality of this income source among the individuals receiving it, $\widehat{Gini}_k$. This pattern was counteracted by an increase in the proportion of household receiving this income source which increased the source inequality component and pushed up the $Gini_k$ of *Wage University Education*.

From these results, it is clear that policies designed to reduce overall inequality in Chile should aim to reduce the inequality between different educational levels, which proved the most important driver of the level of and change in inequality of all income variables analyzed in this paper. It would be informative to study the source of the large income disparities that exist between Chile's least- and most-educated workers. A potential policy to reduce the income gap between these groups involves the increased training and education of low-skilled workers in order to raise their productivity and wages. Another potential policy is to expand the number of and lower the tuition fees at Chile's universities; such policy would increase the supply of university-educated workers, reduce the education wage premium, and thereby generate a Compression effect of the between-group inequality (Knight and Sabot, 1983). Additionally, policies for self-employed workers might be needed. Such policy possibilities include micro-credits targeted at low-income, self-employed workers, and programs to increase the opportunities of women to combine family responsibilities with full-time work.

Another group attracting increased attention among Chilean scholars, but not studied in this paper, is that comprised of wage and salary workers without contracts. This group has grown during the 1990s (Amuedo-Dorantes, 2005), and although some research has addressed the effect of their increased importance, several related issues still require further study. For example, future research might investigate how income inequality for non-contract salary workers differs from that of other labor groups, or why this group increased in size during the 1990s.

9 References

Álvarez, R., and R. Fuentes (2004): "Patrones de Especialización y Crecimiento Sectorial en Chile," *Central Bank of Chile Working papers # 288.*

Amuedo-Dorantes, C. (2005): "Work Contracts and Earnings Inequality: The Case of Chile," *Journal of Development Studies*, 41:4, pp. 589-616.

Athanasopoulos, G., and F. Vahid (2003): "Statistical Inference and Changes in Income Inequality in Australia," *The Economic Record*, 79, pp. 412-424.

Atkinson, A. B. (1997): "Bringing Income Distribution In From the Cold," *The Economic Journal*, 107, pp. 297-321.

Biewen, M. (2002): "Bootstrap Inference for Inequality, Mobility and Poverty Measurement," *Journal of Econometrics*, 1008, pp. 317-342.

Bourguignon, F. (2001): "The Distributional Effect of Growth: Micro vs. Macro Approaches," *Working Paper.*

Bravo, D., D. Contreras, and I. Millán (2002): "The Distributional Impact of Social Expenditure: Chile 1990-1998," Background Paper 2 to *Chile's High Growth Economy, Poverty and Income Distribution 1987-1998*, World Bank Country Study.

Contreras, D. (1996): "Household and Individual Welfare: Evidence from Less Developed Economies," *Ph.D. dissertation*, University of California Los Angeles.

Contreras, D. (2003): "Poverty and Inequality in a Rapid Growth Economy: Chile 1990-1996," *The Journal of Development Studies*, 39(3), pp. 181-200.

Corvo, V., and J. A. Tessada (2003): "Growth and Adjustment in Chile: A Look at the 1990s," *Central Bank of Chile Workings Papers # 204.*

ECLAC (1995): "La Medición de los Ingresos en la Perspectiva de los Estudios de Pobreza: El Caso de la Encuesta CASEN de Chile, años 1987 a 1994," *Unpublished Paper.*

Ferreira, F. H. G., and J. Litchfield (1998): "Calm After the Storms: Income Distribution in Chile, 1987-94," Policy Research Working Paper 1960, World Bank.

Fields, G. S. (2001): "Distribution and Development: A New Look at the Development World," Massachusetts: MIT Press.

Kanbur, R., and N. Lustig (1999): "Why Is Inequality Back in the Agenda?" *Working Paper*, IDB, Washington.

Knight, J. B., and R. H. Sabot (1983): "Educational Expansion and the Kuznets effect," *American Economic Review*, 73:5, pp. 1132-1136.

Le, A. T. (1999): "Empirical Studies of Self-employment," *Journal of Economic Surveys*, 13:4, pp. 381-416.

Lerman, R., and S. Yitzhaki (1994): "Effect of Marginal Changes in Income Source on U.S. Income Inequality," *Public Finance Quarterly*, 22:4, pp. 403-416.

MIDEPLAN (2001): "Pobreza, Crecimiento y Distribución del Ingreso en Chile en los Noventa," *Ministerio de Planificación y Cooperación, Documento # 24.*

Mills, J. A., and S. Zandvakilly (1997): "Statistical Inference via Bootstrapping for Measures of Inequality," *Journal of Applied Econometrics*, 12, pp. 133-150.

Moran, T. P. (2006): "Statistical Inference for Measures of Inequality with a Cross-National Bootstrap Application," *The Economic Record*, 34:3, pp. 296-333.

Palma, A. (2005): "Income Inequality in Chile, Trends and Explanations: A Survey," *Working Paper*, Department of Economics, Göteborg University.

Palma, A. (2006): "Price and Composition Effects in the Rise and Fall of Wages Inequality in Urban Chile," *Working Paper*, Department of Economics, Göteborg University.

Parker, S. C. (1999): "The Inequality of Employment and Self-employment Incomes: A Decomposition Analysis for the U.K.," *Review of Income and Wealth*, 45:2, pp. 263-274.

Pyatt, G., C. Chen, and J. Fei (1980): "The Distribution of Income by Factor Components," *Quarterly Journal of Economics*, 95:3, pp. 451-473.

10 Appendix 1

There are several methods of conducting statistical inference in the analysis of income inequality. One method uses mathematical theory to analytically derive the limiting behavior of the inequality indicator. However, asymptotically driven confidence intervals for inequality measures conducted on fixed samples may not be accurate (Mills and Zandvakilly, 1997). Additionally, asymptotic inference is problematic for the Gini coefficient, and no generally accepted asymptotic procedure exists for conducting hypothesis tests for changes in inequality indicators. Finally, asymptotic theory often generates very complex formulas that quickly become difficult to manage (Moran, 2006). The Jacknife method avoids some of these shortcomings. This method approximates the true sampling through systematic re-sampling by removing one data observation at the time. However, thanks to increased computational power, the Jacknife has been replaced by the Bootstrap method (Moran, 2006).

Bootstrap confidence intervals are computationally inexpensive, easy to calculate, and take into account the bounds of the inequality measures used (Mills and Zandvakilly, 1997). Therefore, this method has become increasingly popular in the analysis of inequality dynamics. More precisely, the Bootstrap provides an estimate of the sampling distribution of the inequality indicator by re-sampling from the empirical sample (Biewen, 2002).

Denote the empirical sample distribution of the (*income*, *weight*) pairs in year s as $\widehat{F}_s$. Following Biewen (2002), I draw B bootstrap samples of n (*income*, *weight*) pairs of observations from $\widehat{F}_s$, i.e.

$$\widehat{F}_s \rightarrow \{(y_1^{*b}, w_1^{*b})_s, (y_2^{*b}, w_2^{*b})_s, ..., (y_n^{*b}, w_n^{*b})_s\} \qquad \text{for} \qquad b = 1, ..., B, \tag{9}$$

and $s = 1990, 1992, ..., 2003$ correspond to the years in which CASEN has been carried out. I then calculate the inequality indicator, I, for each of the Bootstrap sample and denote it as (I_s^{*b}). The standard error of the inequality indicator is given by

$$\widehat{SE_F}(I_s) = SE_{\hat{F}}(I_s^*) = \left[\frac{1}{B-1}\sum_{b=1}^{B}(I_s^{*b} - \overline{I}_s^*)^2\right]^{1/2}, \tag{10}$$

where $\overline{I}_s^* = \frac{1}{B}\sum_{b=1}^{B} I_s^{*b}$.[7] An issue of special concern in this paper is whether inequality changes occurring in the period under consideration are statistically significant. This is easily determined by applying Bootstrap techniques. For instance, if one wants to see whether the inequality change between year s and t is significant, the variable

$$\Delta I_{s,t}^{*b} = I_s^{*b} - I_t^{*b} \qquad b = 1, ..., B. \tag{11}$$

can be calculated. The confidence interval of this variable is calculated by Hall's method

$$Pr(2 * \Delta I_{s,t} - \Delta I_{s,t;HL}^* \leq \Delta I_{s,t} \leq 2 * \Delta I_{s,t} - \Delta I_{s,t;LL}^*) = \frac{100 - 2\alpha}{100} \tag{12}$$

where $\Delta I_{s,t;HL}^*$ and $\Delta I_{s,t;LL}^*$ are the α-th lower and upper percentile of the Bootstrap distribution of the difference in inequality. When zero is not contained in the confidence interval, one can then conclude that a statistically significant change in inequality has taken place. The Bootstrap is used to identify not only significant changes in overall inequality, but also the significance of the change in the within and between-group component of inequality. In this case, the variables I analyze are

[7] An alternative method is to select the elements of the Bootstrap sample with probability in relation to their weights. A preliminary application suggests that at least the same significant changes as those reported in this paper are found with this alternative method.

$$\Delta(Theil - W)^{*b}_{s,t} = (Theil - W)^{*b}_{s} - (Theil - W)^{*b}_{t} \qquad b = 1, ..., B. \tag{13}$$

$$\Delta(Theil - B)^{*b}_{s,t} = (Theil - B)^{*b}_{s} - (Theil - B)^{*b}_{t} \qquad b = 1, ..., B. \tag{14}$$

One previous study applies the Bootstrap method to the analysis of inequality in Chile (Contreras, 1996), but analyzes only the period from 1987 to 1994. Moreover, Contreras calculates the confidence intervals of only the estimators of the level of inequality; changes in inequality are ignored.

11 Appendix 2

Table A1

Changes in the Gini Coefficient ($\Delta Gini$) of Wages and Salaries

		1992	1994	1996	1998	2000	2003
		(1)	(2)	(3)	(4)	(5)	(6)
1990	$\Delta Gini$	0.004	-0.003	-0.001	0.008	0.017	0.003
	(95% CI)	(-0.009;0.017)	(-0.016;0.010)	(-0.016;0.013)	(-0.009;0.024)	(-0.001;0.036)	(-0.013;0.018)
1992	$\Delta Gini$		-0.007	-0.004	0.004	0.014	-0.000
	(95% CI)		(-0.017;0.003)	(-0.018;0.007)	(-0.011;0.018)	(-0.003;0.029)	(-0.015;0.013)
1994	$\Delta Gini$			0.003	0.011	0.021*	0.007
	(95% CI)			(-0.012;0.015)	(-0.004;0.025)	(0.005;0.037)	(-0.0090;0.0196)
1996	$\Delta Gini$				0.008	0.018*	0.004
	(95% CI)				(-0.008;0.025)	(0.001;0.036)	(-0.013;0.020)
1998	$\Delta Gini$					0.010	-0.004
	(95% CI)					(-0.009;0.030)	(-0.022;0.013)
2000	$\Delta Gini$						-0.014
	(95% CI)						(-0.031;0.003)

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. * Represents that the inequality change was significant.

Table A2

Changes in the Theil Index ($\Delta Theil$) of Wages and Salaries

		1992	1994	1996	1998	2000	2003
		(1)	(2)	(3)	(4)	(5)	(6)
1990	$\Delta Theil$	0.002	-0.031	-0.022	0.023	0.026	0.015
	(95% CI)	(-0.035;0.044)	(-0.066;0.010)	(-0.068;0.023)	(-0.065;0.095)	(-0.032;0.083)	(-0.078;0.085)
1992	$\Delta Theil$		-0.032*	-0.023	0.021	0.024	0.013
	(95% CI)		(-0.059;-0.008)	(-0.065;0.010)	(-0.050;0.088)	(-0.033;0.077)	(-0.071;0.077)
1994	$\Delta Theil$			0.009	0.053	0.056*	0.046
	(95% CI)			(-0.026;0.044)	(-0.024;0.118)	(0.005;0.106)	(-0.041;0.104)
1996	$\Delta Theil$				0.044	0.047	0.037
	(95% CI)				(-0.035;0.114)	(-0.011;0.103)	(-0.054;0.103)
1998	$\Delta Theil$					0.003	-0.008
	(95% CI)					(-0.079;0.092)	(-0.116;0.091)
2000	$\Delta Theil$						-0.010
	(95% CI)						(-0.105;0.075)

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. * Represents that the inequality change was significant.

Table A3

Changes in the Log-variance (ΔVar) of Wages and Salaries

		1992	1994	1996	1998	2000	2003
		(1)	(2)	(3)	(4)	(5)	(6)
1990	ΔVar	-0.031*	-0.028	0.005	-0.009	0.047*	0.012
	(95% CI)	(-0.058;-0.005)	(-0.055;0.001)	(-0.025;0.035)	(-0.036;0.019)	(0.002;0.084)	(-0.015;0.040)
1992	ΔVar		0.003	0.036*	0.022	0.078*	-0.043*
	(95% CI)		(-0.018;0.027)	(0.010;0.064)	(-0.001;0.046)	(0.033;0.115)	(0.018;0.069)
1994	ΔVar			0.033*	0.019	0.075*	0.040*
	(95% CI)			(0.007;0.060)	(-0.004;0.043)	(0.035;0.111)	(0.016;0.061)
1996	ΔVar				-0.014	0.042	0.007
	(95% CI)				(-0.041;0.011)	(-0.000;0.080)	(-0.019;0.033)
1998	ΔVar					0.056*	0.021
	(95% CI)					(0.013;0.092)	(-0.006;0.046)
2000	ΔVar						-0.035
	(95% CI)						(-0.071;0.008)

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. * Represents that the inequality change was significant.

Table A4

Theil Index Decomposition of Wages and Salaries

(Theil-W and Theil-B represent within and between-group component)

Variable	1990 (1)	1992 (2)	1994 (3)	1996 (4)	1998 (5)	2000 (6)	2003 (7)
Age							
$Theil - W$	0.412	0.411	0.383	0.390	0.435	0.439	0.427
$Theil - B$	0.026	0.029	0.024	0.025	0.025	0.025	0.026
Gender							
$Theil - W$	0.429	0.431	0.399	0.408	0.454	0.455	0.446
$Theil - B$	0.008	0.009	0.008	0.007	0.023	0.008	0.007
Region							
$Theil - W$	0.419	0.414	0.388	0.388	0.437	0.437	0.431
$Theil - B$	0.018	0.025	0.024	0.027	0.023	0.027	0.023
Education							
$Theil - W$	0.319	0.287	0.278	0.273	0.287	0.274	0.286
$Theil - B$	0.119	0.152	0.129	0.143	0.173	0.190	0.167
Participation							
$Theil - W$	0.434	0.434	0.405	0.409	0.448	0.451	0.440
$Theil - B$	0.004	0.006	0.002	0.005	0.012	0.013	0.013
Occupation							
$Theil - W$	*n.a.*	0.363	0.345	0.337	0.379	0.379	0.378
$Theil - B$	*n.a.*	0.076	0.062	0.080	0.081	0.085	0.075
Sector							
$Theil - W$	*n.a.*	0.434	0.399	0.409	0.453	0.462	0.449
$Theil - B$	*n.a.*	0.006	0.008	0.005	0.007	0.002	0.005

Source: Author's own calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. *n.a.* the decomposition cannot be done for this year since the classification of occupation and sector of 1990 is different from that of the other years.

Table A5

Changes in the Gini Coefficient ($\Delta Gini$) of Self-employment Incomes

		1992	1994	1996	1998	2000	2003
		(1)	(2)	(3)	(4)	(5)	(6)
1990	$\Delta Gini$	-0.004	0.011	-0.018	-0.022*	0.006	0.009
	(95% CI)	(-0.026;0.017)	(-0.011;0.033)	(-0.042;0.008)	(-0.042;-0.001)	(-0.035;0.043)	(-0.015;0.034)
1992	$\Delta Gini$		0.016	-0.013	-0.017	0.011	0.014
	(95% CI)		(-0.006;0.035)	(-0.038;0.012)	(-0.034;0.004)	(-0.029;0.051)	(-0.007;0.038)
1994	$\Delta Gini$			-0.029*	-0.033	-0.005	-0.002
	(95% CI)			(-0.056;-0.005)	(-0.051;0.011)	(-0.047;0.034)	(-0.024;0.020)
1996	$\Delta Gini$				-0.004	0.024	0.027*
	(95% CI)				(-0.026;0.021)	(-0.021;0.065)	(0.002;0.054)
1998	$\Delta Gini$					0.029	0.031*
	(95% CI)					(-0.012;0.067)	(0.010;0.052)
2000	$\Delta Gini$						0.003
	(95% CI)						(-0.036;0.048)

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. * Represents that the inequality change was significant.

Table A6

Changes in the Theil Index ($\Delta Theil$) of Self-employment Incomes

		1992	1994	1996	1998	2000	2003
		(1)	(2)	(3)	(4)	(5)	(6)
1990	$\Delta Theil$	-0.011	0.017	-0.103*	-0.097*	0.002	0.038
	(95% CI)	(-0.092;0.072)	(-0.079;0.102)	(-0.195;-0.019)	(-0.182;-0.019)	(-0.164;0.146)	(-0.065;0.140)
1992	$\Delta Theil$		0.029	-0.092	-0.086*	0.013	0.049
	(95% CI)		(-0.071;0.108)	(-0.175;0.005)	(-0.166;-0.002)	(-0.151;0.155)	(-0.060;0.147)
1994	$\Delta Theil$			-0.121*	-0.115*	-0.016	0.021
	(95% CI)			(-0.219;-0.018)	(-0.196;-0.030)	(-0.173;0.138)	(-0.086;0.124)
1996	$\Delta Theil$				0.006	0.105	0.141*
	(95% CI)				(-0.084;0.099)	(-0.055;0.248)	(0.027;0.254)
1998	$\Delta Theil$					0.099	0.135*
	(95% CI)					(-0.071;0.243)	(0.019;0.229)
2000	$\Delta Theil$						0.036
	(95% CI)						(-0.119;0.202)

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. * Represents that the inequality change was significant.

Table A7

Changes in the Log-variance (ΔVar) of Self-employment Incomes, 1990-2003

		1992	1994	1996	1998	2000	2003
		(1)	(2)	(3)	(4)	(5)	(6)
1990	ΔVar	-0.125*	0.048	-0.050	-0.127*	0.026	0.072
	(95% CI)	(-0.199;-0.046)	(-0.026;0.127)	(-0.139;0.033)	(-0.203;-0.051)	(-0.101;0.136)	(-0.004;0.148)
1992	ΔVar		0.172*	0.075	-0.002	0.151*	0.128*
	(95% CI)		(0.107;0.235)	(-0.001;0.156)	(-0.072;0.065)	(0.024;0.253)	(0.128;0.264)
1994	ΔVar			-0.098*	-0.174*	-0.021	0.025
	(95% CI)			(-0.185;-0.018)	(-0.242;-0.108)	(-0.142;0.090)	(-0.044;0.093)
1996	ΔVar				-0.077	0.076	0.122*
	(95% CI)				(-0.145;0.004)	(-0.051;0.193)	(0.041;0.202)
1998	ΔVar					0.153*	0.198*
	(95% CI)					(0.033;0.265)	(0.129;0.265)
2000	ΔVar						0.046
	(95% CI)						(-0.062;0.178)

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. * Represents that the inequality change was significant.

Table A8

Theil Index Decomposition of Self-employment Incomes

(Theil-W and Theil-B represent within and between-group component)

Variable	1990 (1)	1992 (2)	1994 (3)	1996 (4)	1998 (5)	2000 (6)	2003 (7)
Age							
$Theil - W$	0.801	0.785	0.818	0.701	0.708	0.806	0.831
$Theil - B$	0.023	0.027	0.024	0.022	0.021	0.028	0.026
Gender							
$Theil - W$	0.806	0.800	0.840	0.712	0.715	0.797	0.834
$Theil - B$	0.018	0.011	0.001	0.010	0.014	0.036	0.023
Region							
$Theil - W$	0.811	0.773	0.802	0.673	0.690	0.801	0.788
$Theil - B$	0.013	0.039	0.039	0.050	0.039	0.033	0.069
Education							
$Theil - W$	0.666	0.586	0.677	0.500	0.455	0.528	0.538
$Theil - B$	0.158	0.226	0.165	0.223	0.274	0.306	0.319
Participation							
$Theil - W$	0.798	0.783	0.820	0.686	0.695	0.783	0.795
$Theil - B$	0.027	0.029	0.022	0.036	0.034	0.051	0.062
Occupation							
$Theil - W$	*n.a.*	0.678	0.701	0.580	0.593	0.671	0.704
$Theil - B$	*n.a.*	0.134	0.140	0.142	0.136	0.163	0.153
Sector							
$Theil - W$	*n.a.*	0.794	0.820	0.702	0.724	0.832	0.844
$Theil - B$	*n.a.*	0.017	0.021	0.020	0.005	0.002	0.013

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. *n.a.* the decomposition cannot be done for this year since the classification of occupation and sector of 1990 is different from that of the other years.

Table A9

Inequality and Relative Mean Within Groups of Wages and Salaries and Self-employment Incomes

Variable	Wages and Salaries						Self-employment Incomes					
	$Theil_i$			μ_i/μ			$Theil_i$			μ_i/μ		
	1990	1996	2003	1990	1996	2003	1990	1996	2003	1990	1996	2003
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Age												
15-24	0.20	0.19	0.14	0.62	0.60	0.54	0.73	0.45	0.39	0.48	0.51	0.36
25-34	0.36	0.33	0.32	0.97	0.96	0.96	0.70	0.51	0.66	0.90	0.81	0.85
35-44	0.45	0.47	0.41	1.20	1.19	1.06	0.75	0.70	0.67	1.02	1.08	0.95
45-54	0.53	0.45	0.63	1.26	1.23	1.25	0.78	0.67	0.85	1.16	1.23	1.18
55-64	0.48	0.47	0.50	1.02	1.04	1.17	0.98	0.86	0.94	1.20	0.90	1.08
$\geq$ 65	0.96	0.39	0.60	1.22	0.82	1.07	0.94	1.03	1.38	1.09	1.08	1.25
Gender												
Male	0.38	0.44	0.49	1.09	1.08	1.09	0.84	0.71	0.85	1.11	1.09	1.14
Female	0.30	0.33	0.35	0.81	0.84	0.85	0.67	0.71	0.78	0.69	0.78	0.69
Region												
I	0.38	0.30	0.41	1.21	1.06	0.96	0.75	0.58	0.53	1.34	0.83	0.72
II	0.36	0.32	0.33	1.38	1.35	1.26	0.67	0.73	0.45	1.07	1.25	0.68
III	0.43	0.36	0.31	1.16	1.07	0.89	0.83	0.70	0.39	1.12	1.22	0.53
IV	0.29	0.35	0.37	0.74	0.75	0.84	0.87	0.58	0.69	0.83	0.72	0.68
V	0.44	0.34	0.32	0.89	0.90	0.84	0.85	0.44	0.50	0.85	0.72	0.67
VI	0.34	0.29	0.28	0.83	0.69	0.74	0.70	0.65	0.40	0.77	0.71	0.67
VII	0.30	0.37	0.37	0.68	0.64	0.67	0.94	0.66	0.84	0.92	0.72	0.61
VIII	0.43	0.39	0.37	0.83	0.84	0.80	0.98	0.71	0.75	0.89	0.81	0.73
IX	0.44	0.35	0.35	0.77	0.69	0.77	1.03	0.70	0.94	0.74	0.53	0.62
X	0.36	0.33	0.30	0.76	0.70	0.78	0.94	0.62	0.64	0.97	0.67	0.73
XI	0.35	0.32	0.36	1.05	0.85	1.10	0.53	0.51	0.59	0.93	0.68	0.84
XII	0.37	0.32	0.29	1.16	1.03	1.08	0.54	0.47	0.86	0.82	0.85	1.09
Metropolitan	0.45	0.42	0.51	1.15	1.21	1.20	0.72	0.71	0.88	1.16	1.34	1.44
Education												
$\leq$ 8	0.31	0.18	0.14	0.61	0.52	0.51	0.64	0.40	0.38	0.56	0.46	0.41
9-12	0.24	0.22	0.20	0.83	0.82	0.72	0.82	0.59	0.59	1.09	0.98	0.77
13-16	0.32	0.31	0.45	1.34	1.35	1.28	0.56	0.54	0.53	1.50	1.83	1.42
$\geq$ 17	0.43	0.36	0.32	2.43	2.52	2.58	0.51	0.41	0.56	2.93	3.14	3.92
Participation												
Part-time	0.47	0.51	0.55	0.73	0.70	0.61	0.86	0.73	0.88	0.60	0.52	0.48
Full-time	0.43	0.40	0.43	1.03	1.04	1.06	0.79	0.68	0.78	1.12	1.13	1.21
Occupation												
Blue-collar	*n.a.*	0.21	0.19	*n.a.*	0.64	0.63	*n.a.*	0.42	0.39	*n.a.*	0.54	0.51
White-collar	*n.a.*	0.41	0.47	*n.a.*	1.43	1.40	*n.a.*	0.65	0.83	*n.a.*	1.63	1.62
Sector												
Primary	*n.a.*	0.56	0.50	*n.a.*	0.77	0.78	*n.a.*	0.80	0.81	*n.a.*	0.59	0.63
Industry	*n.a.*	0.37	0.38	*n.a.*	1.01	0.98	*n.a.*	0.80	0.95	*n.a.*	1.03	1.04
Service	*n.a.*	0.40	0.46	*n.a.*	1.06	1.06	*n.a.*	0.66	0.81	*n.a.*	1.10	1.07

Source: Author's calculations from CASEN. *n.a.* the decomposition cannot be done for this year since the classification of occupation and sector of 1990 is different from that of the other years.

Table A10

Changes in the Gini Coefficient ($\Delta Gini$) of Household Income per Adult Equivalent

		1992	1994	1996	1998	2000	2003
		(1)	(2)	(3)	(4)	(5)	(6)
1990	$\Delta Gini$	-0.006	-0.015*	0.002	0.007	0.016	-0.002
	(95% CI)	(-0.019;0.006)	(-0.029;-0.002)	(-0.016;0.018)	(-0.006;0.021)	(-0.008;0.036)	(-0.017;0.013)
1992	$\Delta Gini$		-0.009	0.008	0.014*	0.022*	0.005
	(95% CI)		(-0.023;0.003)	(-0.009;0.024)	(0.000;0.027)	(0.001;0.040)	(-0.019;0.005)
1994	$\Delta Gini$			0.018	0.023*	0.032*	0.014*
	(95% CI)			(-0.000;0.035)	(0.011;0.036)	(0.009;0.054)	(0.001;0.028)
1996	$\Delta Gini$				0.005	0.014	-0.004
	(95% CI)				(-0.011;0.025)	(-0.011;0.037)	(-0.021;0.014)
1998	$\Delta Gini$					0.009	-0.009
	(95% CI)					(-0.013;0.030)	(-0.024;0.004)
2000	$\Delta Gini$						-0.018
	(95% CI)						(-0.037;0.004)

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications.* Represents that the inequality change was significant.

Table A11

Changes in the Theil Index ($\Delta Theil$) of Household Income per Adult Equivalent

		1992	1994	1996	1998	2000	2003
		(1)	(2)	(3)	(4)	(5)	(6)
1990	$\Delta Theil$	-0.011	-0.049	-0.011	0.010	0.049	0.020
	(95% CI)	(-0.055;0.038)	(-0.108;0.003)	(-0.075;0.048)	(-0.044;0.067)	(-0.025;0.124)	(-0.047;0.081)
1992	$\Delta Theil$		-0.037	-0.000	0.022	0.061	0.032
	(95% CI)		(-0.093;0.013)	(-0.060;0.060)	(-0.028;0.070)	(-0.013;0.133)	(-0.032;0.090)
1994	$\Delta Theil$			0.037	0.059*	0.098*	0.069*
	(95% CI)			(-0.031;0.098)	(0.007;0.115)	(0.015;0.183)	(0.005;0.140)
1996	$\Delta Theil$				0.021	0.061	0.031
	(95% CI)				(-0.037;0.088)	(-0.022;0.145)	(-0.040;0.101)
1998	$\Delta Theil$					0.038	0.001
	(95% CI)					(-0.043;0.120)	(-0.055;0.068)
2000	$\Delta Theil$						-0.029
	(95% CI)						(-0.111;0.060)

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. * Represents that the inequality change was significant.

Table A12

Changes in the Log-variance (ΔVar) of Household Income per Adult Equivalent

		1992 (1)	1994 (2)	1996 (3)	1998 (4)	2000 (5)	2003 (6)
1990	ΔVar	-0.079	-0.043*	0.029	0.072*	0.060*	-0.018
	(95% CI)	(-0.115;-0.042)	(-0.081;-0.002)	(-0.024;0.079)	(0.016;0.121)	(0.003;0.114)	(-0.057;0.020)
1992	ΔVar		0.036	0.108*	0.151*	0.139*	0.060*
	(95% CI)		(-0.001;0.069)	(0.061;0.149)	(0.100;0.199)	(0.084;0.186)	(0.027;0.092)
1994	ΔVar			0.072*	0.115*	0.103*	0.025
	(95% CI)			(0.029;0.118)	(0.064;0.164)	(0.046;0.154)	(-0.007;0.061)
1996	ΔVar				0.043	0.031	-0.047
	(95% CI)				(-0.009;0.099)	(-0.037;0.090)	(-0.090;0.004)
1998	ΔVar					-0.012	-0.090*
	(95% CI)					(-0.070;0.046)	(-0.135;-0.041)
2000	ΔVar						-0.078*
	(95% CI)						(-0.127;-0.027)

Source: Author's calculations from CASEN.

Notes: Calculated using 500 Bootstrap replications. * Represents that the inequality change was significant.

Occupational Structure and Earnings Inequality in Chile, 1992-2000

ALEXIS PALMA †‡

Abstract

This paper studies the distributional effects of an occupational change that occurred in the Chilean labor market from 1992 to 2000. During this period the occupational structure shifted towards informal employment (from 9% in 1992 to 15% in 2000), but also towards professional occupations (from 19% in 1992 to 25% in 2000).

Earnings inequality, defined as the variance of log-earnings, increased from 0.771 to 0.828 over the research period. Inequality increased also within employment sectors, specially among self-employed workers. Within occupations the picture is more varied with respect to size and sign of the inequality change.

Both within-group and between-group composition-effects increased total inequality, while within-group change in variance reduced it. However, using the inequality decomposition of Fields and Yoo (2000) to analyze the inequality change within-occupations it is found that even education and the number of hours worked had an increasing effect on overall inequality. All the other variables in the earnings equation, and especially the residuals, had an inequality decreasing effect.

†Department of Economics, Göteborg University, Box 640, SE-405 30 Göteborg, Sweden. Phone: +46 (0)31 786 13 67. E-mail: alexis.palma@economics.gu.se

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1 Introduction

Chile's recent economic development is unprecedented in its modern economic history. After recovering from a severe economic crisis in the early 1980s, most key economic indicators have moved in the right direction since the late-1980s, characterizing the 1990s as a decade of considerable socioeconomic progress. The high growth rate of Gross Domestic Product (GDP) during this decade, on average 6.3%, implied increased labor market opportunities that were translated into a substantial decline of poverty rates, from 40% in 1987 to less than 20% in 2000; an achievement with few counterparts outside East Asia (Gill and Montenegro, 2002). This poverty reduction have moved the attention to inequality issues since income disparities have remained stable at a high level and even increased during the second half of the 1990s (see Table 1). This occurred even though labor market, for most of the decade, has worked in an economic environment propitious for the reduction of inequality with high output growth, declining unemployment, and increasing minimum wage. Nevertheless, the Chilean labor market has been experiencing a shift in the occupational structure towards low paid and low inequality informal employment, but also towards highly paid and high inequality professional occupations, a pattern that has received relatively little attention in the Chilean economic literature.

More specifically, the structure of employment during the 1990s can be characterized by three different facts. First, Amuedo-Dorante (2004) noticed that the share of informal employment increased during the 1990s. In 1992, when growth was over 12% and employment 11% higher than in 1990, formal wage-employment reached its highest levels of the decade, whereas informal wage-employment represented only 10% of the employed labor force (Table 1).

Table 1

Social and Economic Indicators for Chile, 1990-2000

	GDP growth rate	*Employment index*	*Unemployment rate*	*Female participation*	*Formal employment %*	*Informal employment %*	*Self-employment %*	*Earnings inequality all workers*	*Earnings inequality salaried workers*	*Earnings inequality self-employed workers*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
1990	3.7	1.00	7.4	31.7	62.14	12.02	25.84	0.79	0.61	1.20
1992	12.3	1.11	6.2	33.6	64.34	10.38	25.28	0.77	0.59	1.11
1994	5.7	1.16	7.8	35.4	60.64	14.27	25.10	0.77	0.60	1.24
1996	7.4	1.19	5.4	34.5	60.05	15.90	24.05	0.87	0.66	1.28
1998	3.9	1.21	7.2	36.1	59.01	16.82	24.17	0.83	0.63	1.22
2000	5.4	1.17	8.3	35.1	59.41	16.50	24.09	0.89	0.71	1.36

Sources: (1) and (4) Chile Social and Economic Indicators 1960-2000; others are own calculations using CASEN data.

Notes: (1) Growth rate of real Gross Domestic Product; (2) index of total number of employed workers; (3) October-December of each year; (4) October-December of each year as a share of individuals older than 15 years of age; (6) informal workers are defined as wage-earners without a written contract; (7) self-employment includes own-account workers and employers; (8)-(10) Variance of log-monthly earnings.

In 2000, Chile was recovering from the effects of the Asian crisis, consequently, the rate of unemployment increased and considerable problems to create new jobs appeared (Cowan *et al.*, 2005). Informal employment was quite high; formal employment was five percentage points lower than in 1992. Second, there was a major change in the employment structure towards professional occupations and away from unskilled occupations, from 1992 to 2000. Professionals, technicians, and public administrators, who together constituted 19% of the working labor force in 1992, were up to 25% in 2000, whereas unskilled workers fell from over 26% to under 21%. Third, the share of self-employment was notable stable during a period that the economy and the labor market expanded.

Since the population share of occupational groups traditionally found at the bottom and top of the earnings distribution increased in Chile's labor market, and the level of inequality differs across these groups, it is highly likely that the changing employment structure affected the level and structure (within and between-group inequality) of earnings inequality. In recent research, most attention has been devoted to education (Contreras, 2002), unemployment (Beyer, 1997; Larrañaga, 2001; Meller *et al.*, 1996), and minimum wages (World Bank, 1997; Meller *et al.*, 1996) when the labor market has been analyzed in a context of inequality but relatively little research has been devoted to the effect of the increased share of informal employment and the change in the occupational structure on the level and change of inequality. Neither has research shed light on whether the effect that education has on inequality and its change differ across occupations and sectors. This is an important aspect as several studies point out that education is a crucial determinant of the level and change of inequality in Chile. The aim of this study is to elucidate how the aforementioned shift in the occupational structure affected the distribution of earnings during the 1990s.

I also intend to analyze the main individual factors that influenced the choice among twelve labor market groups, and the main variables explaining earnings-inequality within these groups, as well as the changes therein from 1992 to 2000.

In this paper, I applied an inequality-decomposition with within-group and between-group components, and separated the effect of occupational shift in each of these components. To identify which factors are important to explain the inequality-change within occupations, I used the inequality-decomposition of Fields and Yoo (2002). In the Chilean case, a similar decomposition was used by Uthoff (1986) when he analyzed data from Santiago. What distinguishes my analysis is the use of national household data and another sector classification, both of which reflect the occupational change of the 1990s more accurately.

The individuals examined in this paper are males and females between 20 and 65 years of age, classified as non-working or belonging to one of the 11 occupational groups which are defined according to sector of employment (self-employment, informal employment, or formal employment) and occupation (professional, service, skilled blue-collar, and unskilled). This partition generates a totally of twelve labor market groups.[1]

Earnings inequality within-occupations decreased between 1992 and 2000 in the occupations of the formal employment sector. In the informal sector, inequality increased in all occupations, but it declined in three of the four occupational groups of the self-employment sector. My findings suggest that the composition effects and the change in variance of the between-group component had an inequality increasing effect from 1992 to 2000, while the effect of change in variance of the within-group component worked in the opposite direction.

[1]Informal workers were classified only as service, skilled blue-collar, and unskilled.

Using the estimates of the earnings regression and Fields and Yoo (2000) inequality-decomposition, I am able to explain 38%-45% of the inequality of earnings among professional workers in 2000. Education was found to be a key variable for the earnings inequality in these occupations, accounting for 12%-19%. Among unskilled occupations, I am able to explain 22%-39%. Among these workers, education explained a lower share of their earnings inequality (highest among individuals in the formal sector) than did hours of work, which contributed 30% of the earnings inequality among self-employed workers, and 34% among informal workers, but only 6% among their formal counterparts.

The next section introduces the occupational groups used in this study. Section 3 presents the total inequality decomposition, the occupation choice model, and specification of the model. Section 4 gives then a full description of the data. Section 5 reports the results of the estimation of the multinomial logit model, of the earning equations, and of the inequality decomposition; followed by a section that provide a summary and conclusions.

2 Labor Market Taxonomy

The primary concern of the present paper is the distributional effect of the occupational shift observed between 1992 and 2000 in Chile, and the variables that affected how inequality changed within occupations. Since the employment structure changed both across sectors and across occupations, I distinguish twelve labor market groups that allows to analyze the labor market in these two aspects. First, the individuals outside the labor market are those that *Caracterisazión Socioeconómica Nacional* (CASEN) classified as neither unemployed nor working. Then, working individuals are classified into one of three different sectors: self-employed, informal, and formal. The self-employed are those who work for themselves and may hire employees, employers, or may work alone, the so-called own-account workers.

The underlying reason for having the self-employed as a separate group, and not making them a part of the informal sector[2], is that the allocation of paid work between self-employment and wage-employment has attracted increased interest in the international literature in recent years, as self-employment has become more important in industrialized countries' labor markets (Le, 1999). In developing economies, the percentage of self-employed is usually large, yet knowledge about self-employment is still sketchy (Yamada, 1996). In the Chilean case, its importance is reflected by Table 1, which suggests that 24% of the employed workers are found in the self-employment sector. Moreover, in a previous work I found that the socioeconomic structure of the self-employed differs in many aspects from that of wage and salary workers, such as gender and educational composition (Palma, 2006).

To identify informal employment is rather complicated since there is no international consensus about how it should be defined. In several works, its definition is related to the size of the production unit, employment status (self-employment, wage and salary worker), and occupation (professional, non-professional); (Saavedra and Chong, 1999). Since the self-employed are treated separately in this paper, it is only for wage and salary workers that I need a rule of classification. The definition depending on the size of the firm has received criticism, and it has instead been suggested that such a definition should rely on legalistic aspects (Saavedra and Chong, 1999). In line with a previous work on Chilean data (Amuedo-Dorantes, 2004), informal workers are wage-earners lacking a written contract. I nevertheless add the requirement of being a non-professional to exclude doctors, lawyers and accountants

[2]as in Uthoff (1986) and Mizala and Romaguera (1996)

from informal employment. This is thus a flexible employment relationship which the employer can terminate at any time, and where the wage is not subject to any regulation such as minimum wage requirements or dismissal costs.

Formal workers are defined as wage and salary workers with any type of written contract. Self-employed and formal workers were then classified into one of four occupations: professionals, service, skilled blue-collar, unskilled (see table A2 for more detailed information for the occupations included in each group). Informal workers, on the other hand, were only classified into the three less skilled occupations. Figure 1 in the appendix shows the resulting structure of the labor market. A similar occupational classification is found in the classical work of Schmidt and Strauss (1975) analyzing US data. An application on African data is found in De Beyer and Knight (1989), and in Latin America in Brown *et al.* (1999).

Typical occupations among self-employed professionals are managers, accountants, and lawyers running an own business; while among the unskilled self-employed a large group are street sellers and independent peons in farming. In a disaggregated inspection one finds domestic servants, peons in farming, sellers, and drivers among the informal workers. Formal professionals are constituted by teachers, accountants, engineers, doctors, and a great variety of professionals in different sectors of the economy. Domestic servants, cleaners and unskilled hotel workers, porters and guards, and peons in farming are the type of workers found among unskilled formal workers.

3 Model and Specification

3.1 The Inequality Decomposition

Income inequality can be decomposed into income components (e.g. labor income, capital income, transfers) or as here by population subgroups. I measure inequality using the variance of log-earnings, which can be decomposed into the sum of inequality between-group and inequality within-group. For 2000 this inequality measure, σ^2_{2000}, can be written in the following way:

$$\sigma^2_{2000}(\ln E) = \sum_{s=2}^{12} P_{s(2000)} * \sigma^2_{s(2000)} + \sum_{s=2}^{12} P_{s(2000)} * (\ln E_{s(2000)} - \overline{\ln E}_{(2000)})^2 \quad (1)$$

where P_s is the employment share of occupational group s, σ^2_s is the variance of log-earnings within occupation s, $\ln E_s$ is mean log-earnings in that occupation; and $\overline{\ln E}$ is the mean of log earnings of the entire sample. This expression implies that inequality depends on the structure of employment, on the inequality within the occupational groups, and on the distance between the occupation and sample mean of log-earnings. A direct way to quantify the effect of a change in the occupational composition is using the first difference version of Equation (1), which yields;

$$\begin{aligned}\Delta\sigma^2(\ln E) = \sum_{s=2}^{12} \Delta P_s * \sigma^2_{s(1992)} + \sum_{s=2}^{12} P_{s(2000)} * \Delta\sigma^2_s + \\ \sum_{s=2}^{12} \Delta P_s * (\ln E_{s(1992)} - \overline{\ln E}_{(1992)})^2 + \sum_{s=2}^{12} P_{s(2000)} * \Delta(\ln E_s - \overline{\ln E})^2 \quad (2)\end{aligned}$$

where Δ denotes the change between 1992 and 2000. The first term is the within-group inequality-change due to change in occupational shares (the composition effect). The second is the effect of the change in occupational variance which changes the inequality within-group but keeps the employment structure as in 2000. The third term is the composition effect on between-group inequality change, while the last part is attributable to the change in between-group variance.[3] According to Equation (2), inequality increases if workers are moving from low-inequality to high-inequality occupations, or from occupations with mean log-earnings near to the sample mean to occupations with mean of log-earnings far from the sample mean. Moreover, inequality might increase if inequality within occupations grows, or if the difference among occupation mean log-earnings and sample mean log-earnings grows.

Further decompositions are possible, but I focus on the changes in inequality within occupations, $\Delta\sigma_s^2$. I use an inequality decomposition proposed by Fields and Yoo (2000) used previously on Chilean data by Contreras (2002), Contreras (2003), and Amuedo-Dorantes (2005). The appealing characteristic of this inequality decomposition is that it is based in a regression model, such as Mincerian equations, and thereby allows to calculate the contribution of education, age, and gender to the inequality within-group. To perform this decomposition, it is necessary to estimate earnings-generating functions for the different occupational groups. However, a problem that arises at this step is that selectivity bias, when individuals are not randomly assigned to the occupational groups, will generate wrong estimates of the parameters. This problem can be solved by introducing in the earnings-equation an additional variable that contains information on the sample selection into the different occupational groups. This approach is based on the work of Lee (1983) who showed that one can generalize Heckman's two-step method to models with more than two outcomes. First an occupational assignment equation is estimated in order to construct an occupational selection-correction variable

[3] A similar decomposition is found in the work of Uthoff (1986) who used Santiago data for 1969 and 1978 for the informal and formal sectors.

$$\lambda_s^i = -\frac{\phi(\Phi^{-1}(P_s^i))}{P_s^i} \quad \forall\, s \geq 2 \tag{3}$$

where ϕ and Φ represent the standard univariate-normal distribution and density functions, respectively, and P_s^i is the predicted probability of occupation s (the occupational choice model is developed in the next subsection). In the second step, λ is an additional variable in the earnings-equation. The resulting equation is (omitting the index i)

$$\ln E_s = \sum_{p=1}^{n-2} \alpha_s^p * x_s^p + \eta_s * \lambda_s + u_s \quad \forall\, s \geq 2 \tag{4}$$

where $x_s : s$ represents variables that explain earnings, u_s is an error-term, and $E(u_s|s) = 0$[4]. Estimating Equation (4) by OLS generates consistent estimates. Following Fields and Yoo (2000), and using the identity

$$cov\left(\sum_{p=1}^{n} \alpha_s^p * x_s^p, \ln E_s\right) = \sum_{p=1}^{n} cov_s(\alpha^p * x^p, \ln E) \tag{5}$$

and the fact that the left-hand side of this expression is the covariance of $\ln E$ and itself, one obtains:

$$\sigma_s^2(\ln E) = \sum_{p=1}^{n} cov_s(\alpha^p * x^p, \ln E) \quad \forall\, s \geq 2, \tag{6}$$

which is a measure of the inequality of log-earnings in occupation s. Since $\rho_s(\alpha^p * x^p, \ln E) = cov_s(\alpha^p * x^p, \ln E)/\sigma_s(\alpha^p * x^p)\sigma_s(\ln E)$, one can replace the expression for $cov_s(\alpha^p * x^p, \ln E)$ in expression (6) above, and dividing both sides by $\sigma_s^2(\ln E)$, one can obtain the share of earnings-inequality in occupation s that relates to a covariate

[4]In the follwing part we replace $\alpha_s^{n-1} = \eta_s$, $x_s^{n-1} = \lambda_s$, $\alpha_s^n = 1$, and $x_s^n = u_s$.

of the earnings equation. For instance, the share related to ("caused by") education is given by

$$S_s^p \quad = \quad \frac{\alpha_s^p * \sigma_s(x^p) * \rho_s(x^p, \ln E)}{\sigma_s(\ln E)} \quad \forall\, s \geq 2 \tag{7}$$

x^p is years of education, $\sigma_s(x^p)$ is the standard deviation of years of education, $\rho_s(x^p, \ln E)$ is the correlation between x^p and log-earnings; and α_s^p is the parameter of years of education in the earnings-equation, interpreted as the rate of return to education; and $\sigma_s(\ln E)$ is the standard deviation of log-earnings. This expression says that S_s^p, which is called the *relative factor-inequality weight*, is positively related to the rate of return to education and to the inequality in education, and negatively related to the level of log-earnings inequality.

In order to explain the change in inequality for a given occupational group, between 1992 and 2000 we write the change in inequality in the following way

$$\Delta\sigma_s^2(\ln E) = \sum_{p=1}^{n}(S_{s(2000)}^p * \sigma_{s(2000)}^2(\ln E) - S_{s(1992)}^p * \sigma_{s(1992)}^2(\ln E)) \tag{8}$$

3.2 The Occupational Choice Model

Probit models are standard for studying the variables that affect the choice between two sectors, such as formal and informal, while multinomial logit models have been used in models with more than two sectors (e.g, Tiefenthaler, 1994). One important characteristic of the multinomial logit models is the assumption Independence of Irrelevant Alternatives (IIA) through the assumption of an extreme-value distribution of the error-terms of the utility function. This assumption implies that the ratio of the probabilities of two different alternatives is completely independent of other alternatives in the choice set.

Let U_s^i denote the utility experienced by individual i in the occupational group s, made up of a systematic part (V_s^i) and a random part (ε_s^i):

$$U_s^i = V_s^i + \varepsilon_s^i, \quad s = 1, 2, ..., 12. \tag{9}$$

Occupation s will be chosen only if $U_s^i >$ max $U_t^i \ \forall\, s \neq t = 1, 2, ..., 12$. I will assume empirically that

$$V_s^i = \beta_s' Y^i \qquad s = 1, 2, ..., 12. \tag{10}$$

where Y^i is a vector of individual and household characteristics affecting individual i's desirability for sector s. If one assumes that the random part of the utility function (and dropping the superscript i for simplicity) ε_s follows a Extreme-Value distribution, we can define the probability that s is chosen as

$$P_s^i = \frac{\exp[\beta_s' Y^i]}{1 + \sum_{s=2}^{12} \exp[\beta_s' Y^i]} \tag{11}$$

where the parameters associated with nonparticipation, $s = 1$, are set to zero for identification since individuals select only one occupational group and thereby only 11 sets of coefficient are uniquely defined. One can rewrite the log odds as:

$$\ln\left(\frac{P_s^i}{P_1^i}\right) = \ln\left(\frac{\exp[\beta_s' Y^i]}{\exp[\beta_1' Y^i]}\right) = \beta_s' Y^i \tag{12}$$

Thus, the slope coefficients measure the impact of the explanatory variable on the log-odd of being in occupation s relative to nonparticipation. Another model is available to study occupational attainment, namely the ordered probit. However, the multinomial logit have become almost standard in this type of analysis and when the two models have been compared the multinomial logit makes more accurate predictions than the ordered probit (Miller and Volker, 1985).

3.3 Model Specification

Amuedo-Dorantes (2004) summarizes two hypothesis about the determinants of informal employment, one *voluntary* and the other *involuntary*. The hypothesis of voluntary employment suggests that preference for employment in a certain sector is completely based on the comparison of expected earnings in the different sectors; see for instance Tiefenthaler (1994) and Gindling (1991). Therefore, human capital variables, such as education and in work experience should be included in the model.

The involuntary hypothesis sees informal employment as much less attractive than formal employment, attracting only individuals that do not find a job in other sectors. Using this approach, Amuedo-Dorante (2004) found evidence that poverty increased the likelihood of working informally in Chile.

The literature suggests that education, work experience, marital status, and other socio-economic conditions may affect the probability of being self-employed. Education might improve managerial ability, probably useful for the self-employed, but it might also make one more attractive for formal employment. Similarly, with work experience one knows better how markets and productive organization works, which might make one more attractive for self-employment (Le, 1999). On the other hand, self-employment is more risky than formal employment, any factor that increases economic insecurity, such as having children or unemployed individuals in the household, might reduce the likelihood of being self-employed.

As regards the choice among occupations, Boskin (1974) modeled the occupational choice based on training costs and potential earnings. Polachek (1979) modeled occupational choice to explain why one may observe different occupational choices among male and female workers. Women's anticipating pregnancy and child rearing chooses occupations with the smallest earnings-losses because of absences from the labor-market. The result is a female segregation into occupations where skills deterioration from disuse is slow. Brown *et al.* (1980) suggest that individuals from large family induce individuals to seek occupations that provide stable employments. In order to capture the aforementioned effects, in vector Y^i I introduce human capital variables such as years of formal education and experience, several family characteristics as the presence of kids in the household, a dummy variable for household heads, a male dummy, a province dummy, and an unemployment dummy.

I use a Mincerian specification on the earnings-equation according to which one should include years of education, a measure of in work experience and its square, and a measure of hours of work. I use age as a proxy of experience since no information on this variable is available in CASEN in both years. Additional variables included are a dummy for male workers to capture the male premium commonly found in empirical studies, and a provincial dummy to capture eventual spatial earnings disparities (Arellano and Braun, 1999). Finally, I use the log of monthly earnings from the principal occupation as the dependent variable.

4 Data Description

The individuals used in this study are all the inactive and working individuals of age 20-65 included in the survey CASEN of 1992 and 2000. This is a cross-sectional, national representative, household survey conducted in 1987, 1990, 1992, 1994, 1996, 1998, 2000, and in 2003 by MIDEPLAN. CASEN contains data on health, housing, education, employment, and various income variables for 35,948 households in 1992 and for 65,036 in 2000. I study only 1992 and 2000 since 1992 is the earliest year for which the information I need to form the occupational groups is available, and 2000 is the latest year available when this paper was written. Moreover, these two years represent two extreme values of inequality levels in the 1990s. The income variable I analyze is the monthly earnings from the principal occupation in October of the mentioned years.

The aggregated participation rate declined by three percentage points over the research period. Nevertheless, not only was the participation rate of different sizes among males and females, it also moved in different directions, declining from 96% to 85% among male, but increasing from 39% to 43% among females (Table A4). Skilled blue-collar workers formed the largest self-employed and formal occupational group in 1992; unskilled workers comprise the largest informal occupation (Table 2).

Professionals was clearly the non-informal group that increased most, implying that a total of 26% of the sample were professionals in 2000 compared with 19% in 1992. In contrast, service occupations remained almost constant, while skilled blue-collar and unskilled fell. Overall, the informal sector grew by 6 percentage points, from 9% in 1992 to over 15% in 2000, at the same time as the population share of self-employed and formal workers declined. All these changes in the occupational structure were more marked among female than male workers. In fact, the share of informal employment increased from 9% to 13% among male workers but from 10% to 18% among their female counterparts.

There are several potential explanations for the increased share of informal employment during the research period. On the demand-side informal employment (and even self-employment) is a way for employers to avoid dismissal and tax costs as they eliminate these costs when hiring workers without contracts (Amuedo-Dorantes, 2005).

Table 2

Distribution by Occupational Group, 1992 and 2000

Sector	*Occupation*	*1992*		*2000*		*Ratio*
		Total	P_s	*Total*	P_s	*2000/1992*
		(1)	(2)	(3)	(4)	(3)/(1)
Inactive		2250	33.97	2823	36.46	1.25
Active		4374	66.03	4921	63.54	1.13
Total		**6624**	**100**	**7744**	**100**	**1.16**
				Among Employed Workers		
Self-Employment	Professionals	260	5.95	402	8.17	1.55
	Service	194	4.44	157	3.21	0.81
	Skilled blue-collar	483	11.06	503	10.23	1.04
	Unskilled	188	4.32	135	2.75	0.72
Informal Employment	Service	75	1.73	163	3.32	2.17
	Skilled blue-collar	117	2.69	240	4.88	2.05
	Unskilled	218	5.00	338	6.88	1.55
Formal Employment	Professionals	605	13.85	859	17.46	1.42
	Service	681	15.59	816	16.60	1.20
	Skilled blue-collar	860	19.67	770	15.66	0.90
	Unskilled	687	15.71	533	10.84	0.78
Self-Employment		1125	25.77	1197	24.36	1.06
Informal Employment		410	9.42	741	15.08	1.81
Formal Employment		2833	64.82	2978	60.56	1.05
Total		**4368**	**100**	**4916**	**100**	**1.12**

Source: Own calculations from CASEN.

Notes: Columns (1) and (3) are in thousands using weighted data; columns (2) and (4) are percentages.

Since dismissal costs increased in the early 1990s, employers might have employed more informal workers who could be dismissed at no costs. It is also possible that the high growth rate in 1992 (over 12%) fed a growing optimism among employers, who expected growth, and thereby the need for labor, to continue for several years, making informal employment unusually low that year (Table 1). The informal sector in Latin America is traditionally seen as a *buffer* absorbing part of the unemployed when growth declines (Mizala and Romaguera, 1996). Using a definition of informality based on non-professional self-employment, these authors did not find evidence for this hypothesis in Chile. However, Table 1 shows that using the definition of informality based on the contract, between 1990-1994 both unemployment and informal employment followed a U-shaped pattern. This indicates that it is informal wage-employment that might play the role of *buffer*, and not self-employment.

Changes in the minimum wage, which increased by 50% from 1992 to 2000[5] (Ffrench-Davis, 2005), may also have contributed to the increase in informal employment since the minimum wage doesn't apply to this sector.

On the supply side, an explanation to the increased percentage of informal workers is a surge in part-time employment, as indicated by the last column of Table A3 in the Appendix. Part-time work (less than 35 hours per week) was most rare in the occupations of the formal sector, but much more common in the other two sectors, especially among self-employed. However, in the informal sector part-time employment reached its highest levels from 1992 to 2000. This suggests that the preference for part-time work might influence the choice of sector. The most obvious explanation of the surge in part-time employment is the increased participation of women during the 1990s. Already in 1992 the percentage of women was highest in the informal sector, 15%, and it increased most there (to over 21%). Female participation only fell among informal and formal skilled blue-collar, increasing substantially in all other occupational groups. Another explanation is the macroeconomic conditions in 2000, when unemployment was over 8% (compared with 6% in 1992). The worsening conditions in the labor market might not only have induced higher rates of unemployment but also underemployment, especially among those that could reduce the number working hours, such as self-employed and informal workers.

[5]Average wage increased only by 30% over the same period

A possible explanation for the drop in skilled blue-collar and unskilled work and the surge in professional employment is the change in the employment structure among industries. The share of agricultural-hunting-fishing employment decreased from 18% in 1992 to 14% in 2000, and industrial employment from 17% to 14%. On the other hand, the share in financial services increased from 5% to 8%, and in social and communal services from 25% to 28% (Chile Social and Economic Indicators 1960-2000).

Inequality was clearly related to occupation, being highest among professionals in each sector, followed by service occupations, skilled blue-collar, and unskilled. An additional feature is a higher variance of log-earnings among the self-employed occupations than among their counterparts in the other sectors.

Table 3

Variance of Log-Earnings and 95% Confidence Interval by Occupation, 1992 and 2000

Sector	*Occupation*	σ^2_{1992} (1)	σ^2_{2000} (2)	$\Delta\sigma^2$ (3)	95% CI (4)
Self-Employment	Professionals	1.312	1.304	-0.018	(-0.227;0.147)
	Service	0.795	0.849	0.057	(-0.058;0.176)
	Skilled blue-collar	0.765	0.725	-0.040	(-0.106;0.287)
	Unskilled	0.723	0.659	-0.061	(-0.148;0.030)
Informal Employment	Service	0.390	0.540	0.148**	(0.043;0.148)
	Skilled blue-collar	0.356	0.381	0.027	(-0.028;0.091)
	Unskilled	0.350	0.399	0.046**	(0.003;0.091)
Formal Employment	Professionals	0.711	0.631	-0.082**	(-0.132;-0.037)
	Service	0.364	0.363	-0.001	(-0.152;0.092)
	Skilled blue-collar	0.330	0.277	-0.053**	(-0.073;-0.028)
	Unskilled	0.258	0.165	-0.093**	(-0.119;-0.069)
Self-Employment		1.130	1.246	0.108**	(0.003;0.203)
Informal Employment		0.388	0.461	0.071**	(0.031;0.118)
Formal Employment		0.567	0.577	0.010	(-0.039;0.052)
Total		**0.771**	**0.828**	**0.055***	**(-0.001;0.093)**

Source: Own calculations from CASEN.

Notes: Calculated using 100 replications; * significant at 10%; ** significant at 5%; *** significant at 1%.

Changes in earnings inequality were negative within all occupations of the formal sector, and in most of self-employment, while in the informal sector they moved on the opposite direction. Moreover, the 95% confidence intervals of inequality changes indicate that most changes within informal and formal occupations were significant, though by sectors only self-employment and informal sector had significant inequality changes. Figures of occupational inequality differed between males and females, especially in self-employment and professional occupations, but in general they moved in the same direction. However, at sector level only the informal sector reported the same qualitative change.

Another feature of the period I analyze is the clear discrepancy in the growth of average earnings across groups. Changes in average earnings ranged from a small decline for service self-employed workers to a growth over 25% for all occupations of the formal sector (Table 4). In fact, the average earnings of formal unskilled workers increased by more than 35%, but declined by 4% for self-employed with a similar occupation. In consequence, the gap between self-employment and overall earnings declined from 70% to roughly 63% over the research period. For informal workers, on the other hand, this gap changed from -53% to -66% between 1992 and 2000.

In summary, the Chilean labor market experienced an occupational change towards informal and professional occupations, both among male and female workers. But, in general, it is only in the occupations of the formal sector that inequality declined. Those occupations also report the largest earnings growth during the 1990s. In the informal sector, not only inequality increased within occupations, and totally, it also reports lower earnings growth than in the formal sector.

Table 4

Average Earnings in the Different Occupations, 1992 and 2000

Sector	*Occupation*	*1992 Earnings* (1)	*2000 Earnings* (2)	*Ratio 2000/1992* (2)/(1)
Self-Employment	Professionals	865	959	1.11
	Service	269	245	0.91
	Skilled blue-collar	240	244	1.02
	Unskilled	148	142	0.96
Informal Employment	Service	96	117	1.22
	Skilled blue-collar	102	116	1.14
	Unskilled	69	77	1.12
Formal Employment	Professionals	371	507	1.37
	Service	146	186	1.27
	Skilled blue-collar	140	179	1.28
	Unskilled	86	119	1.37
Self-Employment		373	477	1.28
Informal Employment		83	98	1.18
Formal Employment		178	265	1.49
Total		**219**	**292**	**1.33**

Source: Own calculations from CASEN.

Notes: Earnings from principal occupation in thousands of Chilean pesos of 2000.

5 Results

5.1 Estimates of the Occupational Choice Model

This section reports the results of the occupational choice model and of the earnings equations. The analysis was conducted at an aggregated level and not separately for males and females, since my ambition is to explain aggregated inequality. Moreover, due to the large number of occupational groups, an analysis by gender would leave a too small number of individuals in some groups.

Tables 5-6 (below) present the estimates from the multinomial logit model of occupational attainment. The omitted group is non-participation. More years of education have a strong positive effect on participation in professional occupations,

Table 5

Maximum-Likelihood Estimates of the Multinomial Logit Model of Occupational Attainment 1992

Variables	$\ln(P_2/P_1)$	$\ln(P_3/P_1)$	$\ln(P_4/P_1)$	$\ln(P_5/P_1)$	$\ln(P_6/P_1)$	$\ln(P_7/P_1)$	$\ln(P_8/P_1)$	$\ln(P_9/P_1)$	$\ln(P_{10}/P_1)$	$\ln(P_{11}/P_1)$	$\ln(P_{12}/P_1)$
Constant	-8.5072***	-7.1051***	-6.8259***	-4.7970***	-1.8798***	-4.1957***	-2.0989***	-8.1549***	-3.1570***	-3.8987***	-1.8882***
	(19.80)	(17.02)	(22.82)	(11.95)	(3.36)	(8.26)	(6.27)	(24.56)	(12.68)	(15.19)	(8.08)
Ed2	1.0649***	0.5923***	0.0789	-0.6881***	0.7812***	-0.0596	-1.0186***	2.3340***	1.5258***	0.3020***	-0.3696***
	(13.52)	(8.37)	(1.48)	(9.13)	(6.89)	(0.71)	(15.00)	(21.83)	(28.05)	(6.42)	(8.18)
Ed3	2.3866***	1.0277***	0.3022***	-1.0223***	1.3785***	-0.1162	-1.6433***	4.6696***	2.5600***	0.5248***	-0.9454***
	(23.66)	(8.96)	(3.00)	(5.07)	(9.02)	(0.71)	(8.48)	(41.12)	(36.30)	(6.66)	(9.10)
Ed4	3.4057***	0.7059***	-0.0927	-1.9211***	0.5896	-0.9530**	-2.7959***	5.9399***	1.7205***	-0.2054	-2.2899***
	(26.80)	(3.35)	(0.49)	(3.15)	(1.49)	(2.53)	(5.12)	(43.74)	(12.16)	(1.18)	(6.71)
Age	0.2445***	0.2199***	0.1966***	0.1074***	-0.0358	0.0526**	0.0209	0.2236***	0.0995***	0.1236***	0.0612***
	(12.39)	(11.06)	(13.88)	(5.71)	(1.20)	(2.20)	(1.24)	(13.77)	(7.62)	(9.98)	(5.20)
Age^2	-0.0287***	-0.0266***	-0.0247***	-0.0172***	-0.0042	-0.0141***	-0.0097***	-0.0296***	-0.0188***	-0.0213***	-0.0139***
	(12.38)	(11.38)	(14.72)	(7.62)	(1.10)	(4.72)	(4.62)	(14.63)	(11.22)	(14.09)	(9.56)
Male	2.2674***	1.6666***	3.6652***	3.1753***	1.2019***	3.6307***	2.5156***	1.6911***	1.3255***	3.5554***	2.6141***
	(17.20)	(12.70)	(33.50)	(21.86)	(6.99)	(17.67)	(21.61)	(16.18)	(15.13)	(35.43)	(30.77)
Head	1.3064***	1.3375***	1.3478***	1.9468***	1.3802***	1.3725***	1.7280***	1.2717***	1.2301***	1.4427***	1.6036***
	(10.55)	(12.25)	(11.41)	(13.80)	(7.37)	(4.43)	(14.56)	(12.00)	(14.74)	(11.62)	(20.01)
Head*Male	0.9868***	0.4541***	0.3870***	-0.3230*	0.4385*	0.4075	-0.3881***	1.0074***	0.8374***	0.6204***	0.0874
	(5.83)	(2.80)	(2.65)	(1.90)	(1.77)	(1.23)	(2.62)	(6.90)	(6.71)	(4.24)	(0.77)
Kid	-0.7556***	-0.5918***	-0.6407***	-0.3906***	-1.1208***	-0.9174***	-0.5926***	-0.8769***	-1.0262***	-0.8590***	-0.6765***
	(7.20)	(6.69)	(6.18)	(3.01)	(8.55)	(4.20)	(6.05)	(12.44)	(19.01)	(9.20)	(10.55)
Kid*Male	0.6798***	0.7284***	0.8755***	0.6867***	1.2847***	1.2158***	0.9725***	1.0068***	1.1079***	1.1030***	0.9685***
	(4.78)	(5.17)	(6.64)	(4.25)	(6.14)	(5.03)	(7.23)	(8.42)	(10.46)	(8.99)	(9.42)
Province	-0.5771***	-0.5058***	-0.2230***	-0.2428***	-0.4062***	-0.5387***	0.0537	-0.5810***	-0.4657***	-0.7530***	-0.3187***
	(9.27)	(7.67)	(4.31)	(3.43)	(4.03)	(6.55)	(0.82)	(11.26)	(10.89)	(17.23)	(7.23)
Unemployment	-0.2569*	0.0496	0.0054	0.2088*	0.4459***	0.4474***	0.4477***	0.0860	0.3855***	0.4285***	0.2897***
	(1.82)	(0.40)	(0.06)	(1.67)	(2.70)	(3.18)	(4.40)	(0.80)	(5.03)	(5.65)	(3.73)
N (Unweighted)		69743	Prob> χ^2=0.0000								
Pseudo R-squared		0.2327									
Log-likelihood		-111196									

Notes: Robust z-statistics in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 6

Maximum-Likelihood Estimates of the Multinomial Logit Model of Occupational Attainment 2000

Variables	$\ln(P_2/P_1)$	$\ln(P_3/P_1)$	$\ln(P_4/P_1)$	$\ln(P_5/P_1)$	$\ln(P_6/P_1)$	$\ln(P_7/P_1)$	$\ln(P_8/P_1)$	$\ln(P_9/P_1)$	$\ln(P_{10}/P_1)$	$\ln(P_{11}/P_1)$	$\ln(P_{12}/P_1)$
Constant	-11.9946***	-10.2086***	-10.9993***	-8.6947***	-4.3305***	-8.3227***	-6.0394***	-11.9446***	-6.0960***	-9.3600***	-6.5921***
	(26.24)	(19.60)	(30.72)	(14.20)	(9.97)	(19.77)	(21.30)	(32.27)	(23.80)	(31.05)	(23.01)
Ed2	0.8915***	0.5861***	0.2966***	-0.4657***	0.9728***	0.0850	-0.6037***	2.8519***	1.7084***	0.5431***	-0.1105**
	(14.19)	(6.96)	(5.29)	(5.34)	(10.66)	(1.16)	(11.18)	(21.36)	(26.92)	(11.27)	(2.34)
Ed3	1.3401***	0.2442*	-0.4196***	-2.1531***	0.7163***	-1.1994***	-2.4914***	4.4151***	1.6016***	-0.4063***	-1.8742***
	(14.60)	(1.83)	(2.91)	(9.59)	(6.08)	(7.89)	(16.89)	(30.96)	(21.31)	(5.16)	(17.28)
Ed4	2.5896***	0.0878	-1.4565***	-3.5059***	0.5104	-2.4501***	-3.7781***	5.9085***	0.7172***	-1.7590***	-2.7444***
	(19.72)	(0.27)	(7.17)	(4.79)	(1.54)	(6.43)	(6.98)	(38.49)	(4.74)	(8.11)	(7.22)
Age	0.4236***	0.3525***	0.3926***	0.2913***	0.0979***	0.2620***	0.2529***	0.3704***	0.2469***	0.3477***	0.2740***
	(21.12)	(14.79)	(23.39)	(10.15)	(4.37)	(12.30)	(18.40)	(20.64)	(17.74)	(25.43)	(19.86)
Age^2	-0.0480***	-0.0410***	-0.0467***	-0.0374***	-0.0172***	-0.0371***	-0.0356***	-0.0451***	-0.0355***	-0.0467***	-0.0381***
	(21.21)	(15.08)	(23.73)	(11.42)	(6.03)	(13.56)	(21.29)	(20.45)	(20.09)	(28.07)	(22.86)
Male	0.8065***	0.4704***	2.4608***	1.9188***	-0.1131	2.7877***	0.9421***	0.3917***	-0.0145	2.8115***	1.5206***
	(6.72)	(2.92)	(23.23)	(11.23)	(0.80)	(19.79)	(11.76)	(3.32)	(0.17)	(25.94)	(18.11)
Head	1.2547***	1.3671***	1.3651***	1.6942***	1.1014***	0.8715***	1.3878***	1.2108***	1.1637***	1.4187***	1.5482***
	(11.88)	(11.20)	(6.58)	(11.10)	(8.30)	(4.63)	(16.41)	(11.87)	(12.66)	(11.00)	(18.07)
Head*Male	0.8182***	-0.1282	0.1014	-0.7276***	0.0649	0.6326***	-0.3656***	0.7517***	0.6719***	0.4416***	-0.0560
	(5.62)	(0.71)	(0.48)	(3.94)	(0.34)	(3.10)	(3.39)	(5.13)	(5.81)	(3.11)	(0.52)
Kid	-0.5266***	-0.1715*	-0.5459***	-0.0155	-0.3538***	-0.3215**	-0.2582***	-0.4022***	-0.6854***	-0.3741***	-0.5130***
	(5.79)	(1.70)	(4.00)	(0.11)	(3.67)	(2.16)	(3.74)	(5.24)	(9.76)	(3.46)	(7.12)
Kid*Male	0.9891***	0.8548***	1.1215***	0.6617***	1.0945***	1.0069***	0.8893***	0.9331***	1.2942***	1.0645***	1.1737***
	(7.91)	(5.24)	(7.42)	(3.71)	(6.61)	(5.98)	(9.22)	(7.45)	(13.10)	(8.62)	(12.37)
Province	-0.3766***	-0.6136***	-0.0540	-0.3322***	-0.4294***	-0.1595**	-0.1247**	-0.4300***	-0.5518***	-0.3049***	-0.2321***
	(5.72)	(7.66)	(0.83)	(4.10)	(5.58)	(2.27)	(2.34)	(6.59)	(11.58)	(6.08)	(4.75)
Unemployment	-0.4771***	-0.1326	-0.0767	0.0049	0.1376	0.2538***	0.2605***	-0.0235	0.1334**	0.0927	0.1334*
	(4.55)	(0.97)	(1.03)	(0.05)	(1.33)	(2.99)	(3.59)	(0.29)	(2.20)	(1.43)	(1.92)
N (Unweighted)		126650	Prob> χ^2=0.0000								
Pseudo R-squared		0.2081									
Log-likelihood		-207634									

Notes: Robust z-statistics in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

and this effect was stronger for formal workers than for other individuals. Age had a positive and significant effect for most occupations, but the effect was stronger for self-employed, while the coefficient on age squared was negative in all cases and significant in most of them. The coefficient of the male variable is of a larger size than the coefficient of other variables, (except education), and especially for low-skill occupations, thus indicating that men are more likely to choose blue-collar and unskilled occupations over nonparticipation. Being a household head has a positive effect in labor market participation for males and females.

Expectedly, having a kid younger than 10 years has different effects for males and females, being positive for the former and negative for the later group of workers. Living in a region other than Santiago is associated with a lower probability of participation in all occupational groups. Having unemployed members in the household has a significant and negative effect on self-employment, but a positive and significant effect on all occupations except for professional self-employed, for which the effect is negative and significant in both years. Further information about the model is reported in the bottom of the tables. Pseudo-R^2 (=1-L_N/L_0, where L_N is the value of the likelihood-function when all covariates are included, and L_0 is the value when only a constant) for 1992 was 0.2327 but decreased to 0.2081 in 2000. The likelihood ratio test did not reject the joint hypothesis that the parameters were all equal to zero. I also performed Hausman tests to verify if the IIA assumption is fulfilled. All the results supported the assumption. Furthermore, I also tested for the equality of slope parameters for all pairs of occupations. The null hypothesis was clearly rejected at the 1% level in all cases.

5.2 Estimates of the Earnings Equations

The parameters have the expected sign in all occupations. Years of education were significant and positive in all the regressions. Their magnitude was almost identical for self-employed and formal professionals, whereas for the other occupations they were clearly lower among informal workers. The coefficients on Age are all positive and almost all significantly so, and those on Age^2 are generally negative, the majority being significant. The male wage-premium in 1992 was the highest for self-employed skilled blue-collar workers and formal professionals. In 2000, it was the highest for skilled blue-collar workers in all three sectors. This might reflect productivity differences between male and female workers or gender discrimination in occupations where few females work.

Residence in provinces had a negative effect in most cases, larger for self-employed and informal skilled and unskilled workers and formal professionals. The lower average earnings outside Santiago may reflect a compensation for a possibly lower quality of life in the city, for instance because of air pollution (Arellano and Braun, 1999), but might also reflect a higher marginal labor productivity in Santiago. The results here cast doubt on the first suggestion, as compensation for a lower quality of life should not differ substantially across occupations, whereas marginal productivity differences might. The working-hours-elasticity of earnings, that is the effect of the log of hours worked, is less than unity in all cases (and highly significant in most), indicating that hourly-earnings declined as the number of working hours increases.[6]

[6]This result was also found by Riveros (1990) and Arellano and Braun (1999).

Table 7

Estimates of Earnings-Equations by Sectoral Occupation, 1992

(Dependent Variable: Log of Monthly Earnings)

Explanatory	*Self-Employment*				*Informal Employment*			*Formal Employment*			
Variables	*Professional*	*Service*	*Skilled blue − collar*	*Unskilled*	*Service*	*Skilled blue − collar*	*Unskilled*	*Professional*	*Service*	*Skilled blue − collar*	*Unskilled*
Constant	7.784***	8.553***	6.959***	7.025***	7.539***	8.106***	8.117***	6.269***	8.791***	8.192***	8.679***
	(8.758)	(10.874)	(14.314)	(17.024)	(12.390)	(13.771)	(26.733)	(14.618)	(27.932)	(28.007)	(37.434)
Education	0.117***	0.063***	0.067***	0.035***	0.043***	0.030***	0.011**	0.101***	0.081***	0.052***	0.023***
	(8.808)	(8.370)	(15.241)	(3.968)	(4.467)	(4.194)	(2.147)	(8.963)	(16.519)	(20.853)	(8.905)
Age	0.076***	0.022	0.055***	0.058***	0.033**	0.020*	0.005	0.062***	0.051***	0.056***	0.013***
	(3.937)	(1.135)	(5.886)	(4.908)	(2.111)	(1.872)	(0.672)	(7.376)	(8.930)	(10.940)	(3.371)
Age^2	-0.007***	-0.002	-0.005***	-0.007***	-0.004**	-0.002*	-0.001	-0.006***	-0.005***	-0.006***	-0.001***
	(3.408)	(0.797)	(5.113)	(4.737)	(2.081)	(1.695)	(0.742)	(5.333)	(6.424)	(8.338)	(2.644)
Male	0.200**	0.340***	0.629***	0.432***	0.221***	0.260	0.314***	0.445***	0.213***	0.384***	0.346***
	(2.476)	(7.161)	(5.397)	(5.719)	(3.882)	(1.569)	(10.505)	(17.186)	(11.360)	(5.051)	(16.156)
Provincial resident	-0.148***	-0.134***	-0.244***	-0.280***	-0.044	-0.219***	-0.301***	-0.355***	-0.161***	-0.114***	-0.166***
	(2.938)	(2.619)	(7.545)	(5.661)	(0.807)	(4.975)	(9.720)	(14.882)	(8.771)	(5.374)	(11.026)
lnHour	0.467***	0.471***	0.499***	0.483***	0.511***	0.428***	0.494***	0.603***	0.143***	0.259***	0.329***
	(6.939)	(7.882)	(14.459)	(11.159)	(5.715)	(6.055)	(12.200)	(11.469)	(2.786)	(6.111)	(8.151)
λ	0.275	0.095	-0.103	-0.264**	0.031	-0.100	-0.081	0.008	-0.063	-0.100	-0.150***
	(1.600)	(0.523)	(0.846)	(2.028)	(0.189)	(0.494)	(1.353)	(0.153)	(1.503)	(1.361)	(3.816)
N (Unweighted)	2193	1804	5471	2579	681	1250	3024	4800	5514	8302	7943
R-squared	0.320	0.190	0.247	0.237	0.211	0.128	0.209	0.370	0.220	0.161	0.166
F-test	97.751	30.022	148.769	49.952	11.688	20.399	61.711	237.188	124.521	127.048	109.757

Notes: t-statistics in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 8

Estimates of Earnings-Equations by Sectoral Occupation, 2000

(Dependent Variable: Log of Monthly Earnings)

Explanatory	*Self-Employment*				*Informal Employment*			*Formal Employment*			
Variables	*Professional*	*Service*	*Skilled*	*Unskilled*	*Service*	*Skilled*	*Unskilled*	*Professional*	*Service*	*Skilled*	*Unskilled*
			blue − collar			*blue − collar*				*blue − collar*	
Constant	10.640***	6.779***	7.734***	7.298***	6.791***	7.454***	7.815***	7.817***	6.930***	8.428***	8.891***
	(13.730)	(6.344)	(16.643)	(13.950)	(13.711)	(18.736)	(32.432)	(16.597)	(6.359)	(32.390)	(37.274)
Education	0.095***	0.080***	0.074***	0.039***	0.070***	0.032***	0.029***	0.092***	0.088***	0.063***	0.025***
	(9.098)	(8.801)	(15.807)	(4.490)	(8.712)	(5.151)	(6.686)	(6.637)	(14.054)	(25.436)	(9.733)
Age	0.051***	0.057**	0.051***	0.041**	0.022*	0.007	0.016**	0.035***	0.032***	0.049***	0.017***
	(2.691)	(2.132)	(4.730)	(2.199)	(1.785)	(0.674)	(2.313)	(3.119)	(4.454)	(8.514)	(3.625)
Age^2	-0.006***	-0.006*	-0.005***	-0.004*	-0.002	-0.000	-0.001*	-0.002*	-0.003***	-0.005***	-0.002***
	(2.967)	(1.936)	(4.232)	(1.948)	(1.472)	(0.213)	(1.755)	(1.682)	(2.998)	(6.553)	(2.837)
Male	0.257***	0.244***	0.459***	0.284***	0.044	0.465***	0.089***	0.304***	0.182***	0.457***	0.240***
	(4.275)	(3.830)	(4.480)	(4.295)	(0.622)	(4.276)	(4.157)	(10.371)	(9.199)	(8.746)	(8.371)
Provincial resident	-0.205***	-0.100	-0.230***	-0.129***	-0.245***	-0.243***	-0.271***	-0.253***	-0.152***	-0.091***	-0.200***
	(2.895)	(1.564)	(7.884)	(2.768)	(4.841)	(7.633)	(10.926)	(8.836)	(7.321)	(6.135)	(13.672)
lnHour	0.352***	0.506***	0.448***	0.513***	0.545***	0.540***	0.553***	0.530***	0.583***	0.239***	0.339***
	(8.336)	(14.204)	(18.281)	(17.441)	(12.604)	(12.226)	(25.085)	(12.644)	(2.939)	(7.747)	(10.407)
λ	0.868***	-0.219	0.043	-0.205	-0.341*	-0.231**	-0.048	0.186***	-0.110*	-0.114***	-0.160***
	(4.830)	(0.927)	(0.397)	(1.463)	(1.675)	(2.032)	(0.871)	(3.167)	(1.868)	(2.589)	(2.755)
N (Unweighted)	4772	1833	9179	1981	1775	3851	6604	7379	8779	12543	10170
R-squared	0.450	0.336	0.346	0.391	0.352	0.336	0.420	0.386	0.331	0.236	0.227
F-test	133.200	48.141	211.743	63.994	44.131	66.585	135.381	175.663	134.408	171.533	95.620

Notes: t-statistics in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

The parameter for the occupation selection variable, λ_s, was significant in both years for only unskilled formal workers but in the year 2000 the number of occupations with significant selection variable increased to 7. For instance, self-employed and formal professionals had a positive and significant selection parameter which, together with average of the selection variable, indicates that individuals in these occupation categories, on average, have higher earnings than an individual with average characteristics drawn at random from the population would expect to earn, if selected into that labor market group.

5.3 How Much Inequality Can Be Explained?

The results of the inequality-decomposition are reported in Table 9 (below). At first glance, the most striking feature is a clear inequality increasing effect of the composition change, both in the within and in the between-group component. The change in variance, on the other hand, was of distinct sign on the former and later components of the decomposition. Thus, if the occupational structure had remained as it was in 1992, we would have observed a small decline in inequality, but the occupational shift occurred in the period under consideration generated an overall increase.

There are however clear discrepancies in the contribution of the different sectors to the components of the decomposition. The informal sector contributed most to the within-group composition-effect since the percentage employed in the occupations of this sector increased over the research period; while the largest within-group change in variance is found in the formal sector, as a result of the inequality decline in all occupations in this sector. Self-employment had the largest between-group components; on the one hand, because the share of professional occupations on total employment had a large increment in this sector, on the other hand, because self-employment average earnings did not increase as fast as in the formal sector. These changes almost canceled each other out, implying that this sector had a very small total effect. The informal sector had by far the greatest overall effect, a result of the large positive contribution of all components. Accordingly, not only the composition effects contributed the inequality increase, but also to the slow earnings growth in this sector compared to the other two, which widened the between-group component of inequality.

Table 9

Total Decomposition of Changes in Earnings-Inequality, 1992-2000

	Within-group		*Between-group*		***Total***
	Composition-effect	*Change in variance*	*Composition-effect*	*Change in variance*	***contribution***
Self-Employment	0.16	-0.48	2.97	-3.06	**-0.40**
Informal Employment	2.06	0.96	2.06	3.39	**8.47**
Formal Employment	0.35	-3.25	0.06	0.97	**-1.87**
Total contribution	**2.58**	**-2.77**	**5.09**	**1.30**	**6.20**
		Decomposition Within-sectors			
Self-Employment	5.82	-1.96	5.17	1.81	**10.85**
Informal Employment	0.17	6.35	0.05	0.76	**7.32**
Formal Employment	3.41	-5.37	2.16	0.59	**0.80**

Source: Own calculations from CASEN.

Notes: All numbers have been multiplied by 100.

I also calculated the decomposition for males and females separately to see if the results are different from those reported in Table 9. As it was the case when the aggregate group was analyzed, the within-group composition effects were inequality increasing and of similar magnitude for males and females, while the effect of the change in variance of this component was much more important for males (Table A7). Furthermore, when the gender groups were analyzed separately, the sector that contributed most to the composition effect was the informal one, but the contribution was considerable larger for female workers. Between-group components are more or less similar (sign and size) when they are analyzed at an aggregate level and separately by gender. The only exception was perhaps the contribution of the formal sector to the change in variance, which was small and negative for males but large and positive for females. Overall sector figures, last column of Table A7, suggest that while the contributions of self-employment and of the formal sector were of different size across gender, the contribution of the informal sector was large and positive for both gender groups.

After having analyzed the different components of the decomposition, I now focus on what explains the within-group change in variance using the decomposition of Fields and Yoo. The first step is the inspection of the *factor inequality weights* of the variables in the earnings-equation found in Tables A8-A9. In 1992 the contribution of education ranged from less than 1% among informal unskilled workers to 19% among self-employed professionals. In general years of education contributed more to earnings-inequality among professional, service, and skilled blue-collar workers than among unskilled workers, in 2000 as well as in 1992. These results partially corroborate previous studies that used the same inequality decomposition and which found that education was the key determinant of inequality in Chile, accounting for 13-30% depending on sample used (Contreras, 2002; Contreras, 2003; Amuedo-Dorantes, 2005); however, that is not the case here for unskilled workers or for some other non-professionals for whom another variable was more important. In 1992, log of hours-worked explained 9% among self-employed unskilled workers, and 8% among the informal unskilled (but less than 2% among the formal unskilled). In 2000, these contributions had gone up to 25% (self-employed unskilled) and 27% (informal unskilled). In fact, for most non-formal non-professionals in 2000 it contributed more than education. In contrast to this, in the formal sector, the contribution of log of hours-worked was low and remained low over the research period. These results reflect the fact that part-time employment increased in all occupation but from vary small numbers in the formal sector.

Using these factor inequality weights, one is able to decompose the within-group change in variance to evaluate the contribution of the different explanatory variables of the wage equations. Education, log of hours-worked, and the occupation selection variable increased the within-occupation variance for each sector and overall, while all the other variables (and the residual) were negative overall.

Table 10

Contribution of Variables to Within-Group Change in Variance by Sector, 1992-2000

	Education	*Age*	*Male*	*Province*	*lnHour*	λ	*Residual*	***Total contribution***
Self-employment	0.18	-0.11	0.00	0.10	1.11	1.42	-3.12	**-0.43**
Informal employment	0.12	0.02	-0.11	0.11	1.25	0.03	-0.43	**0.99**
Formal employment	0.07	-0.39	-0.65	-0.27	0.99	0.71	-3.70	**-3.24**
Total contribution	**0.37**	**-0.49**	**-0.75**	**-0.07**	**3.35**	**2.16**	**-7.26**	**-2.69**
				Decomposition Within-sectors				
Self-employment	0.73	-0.46	0.00	0.40	4.56	5.83	-12.82	**-1.76**
Informal employment	0.83	0.12	-0.72	0.70	8.31	0.18	-2.85	**6.57**
Formal employment	0.12	-0.65	-1.07	-0.45	1.63	1.18	-6.11	**-5.36**

Source: Own calculations from CASEN.

Notes: All numbers have been multiplied by 100.

Education had its largest effect among the self-employed, as did the selectivity-variable; *lnHour* had its largest effects in the informal sector. Behind the small contribution of education in the formal sector, there are discrepancies in the contribution of the different occupations in this sector, being inequality decreasing for professional occupations but inequality increasing for skilled blue-collar ones. It is understandable that the log of hour-worked had a small contribution to the change in inequality in the formal sector, since it is more difficult to respond to change in macroeconomic conditions through a change in hours of work in this sector, as it is in the other two. This in turn implies that the formal sector respond mainly by increasing unemployment when growth declines. The large inequality-decreasing effect of residuals in the formal sector deserves an explanation. The most tempting is the fast growing minimum wage in the early 1990s and between 1998 and 2000. The minimum wage reduced the dispersion of earnings in the lower part of the distribution, especially in non-professional formal occupations; more and more workers that before had earnings below the new minimum were then granted a higher income.

6 Summary and Conclusions

The present paper studies the choices among twelve occupational groups in Chilean labor market, the distribution of earnings within them, and the effect of a shift in the occupational structure on the inequality-change between 1992 and 2000. Employment in professional occupations and in the informal sector increased during the period, while self-employed and formal unskilled work declined. I explain this pattern by an unusually high GDP-growth rate in 1992, a higher percentage of part-time work, higher female labor-force participation, and the change in employment across industries.

Analyzing the dispersion of earnings by occupations, I found that on the one hand, inequality was substantially higher among professionals than among skilled blue-collar and unskilled workers; on the other hand, inequality was lower among formal unskilled workers and in the formal sector in general. I also found also a clear tendency to lower inequality in the occupations of the formal employment sector, but the pattern was less clear for the other two sectors. These results might be explained by the considerable lower dispersion of hours worked in the the formal sector and by the rapid increase of the minimum wage, which is only applicable in this sector, increased rapidly at the beginning and at the end of the 1990s, decreasing the dispersion of earnings in the lower part of the formal earnings distribution.

The decomposition of the inequality change between 1992 and 2000 indicates that the composition effects and the between-group change in variance increased inequality, whereas the effect of change in variance in the within-group component worked in the opposite direction. Thus, if the occupational structure of 1992 had remained unchanged during the whole decade, one should have seen a drop in inequality, but since the occupational change observed induced a higher inequality, the total effect was an inequality increase between 1992 and 2000. These results reflect the fact that the within-sectors employment structure shifted towards professional occupations and that employment increased in all occupations of the informal sector, hence shifting the structure of employment towards high-inequality occupations. On the other hand, as self- and informal-employment earnings did not increase as much as in the formal sector, self-employment had a negative impact in the between-group change in variance, and the informal sector had an effect of opposite sign. I also found discrepancies in the contribution of different gender groups and sectors. For instance, while the contributions of the change of composition were of similar size for males and females, the within-group change in variance was more important for males than for female workers. Moreover, overall sector figures suggest that while the contributions of self-employment and formal sector were of different sizes across gender, but the contribution of the informal sector was large and positive for both gender groups.

The within-group change in variance was further investigated using a method suggested by Fields and Yoo (2000). My results indicate that the inequality-increasing composition-effect was accompanied by an inequality-increasing effect of education, hour-worked, and selectivity, which reduced the inequality-decreasing effect of the within-group change in variance.

I believe my work has not only shed light on the issues that we where initial investigating, it has also raised other questions and ideas for future research. The direct and indirect cost of education is often neglected in studies on the effect of education on inequality. Arellano and Braun (1999) found that, taking these costs into account, an additional year of primary education increased average earnings by 14% and by 16.5% if the additional year represents university education. If cost is not taken into account, the differences across educational levels are considerably larger. Neglecting the direct and indirect costs of education may exacerbate the effect of education on inequality. This is an interesting issue for future research.

References

Amuedo-Dorantes, C. (2004): "Determinants and Poverty Implications of Informal Sector Work in Chile," *Economic Development and Cultural Change*, 52:2, pp. 347-68.

Amuedo-Dorantes, C. (2005): "Work Contracts and Earnings Inequality: The Case of Chile," *Journal of Development Studies*, 41:4, pp. 589-616.

Arellano, M. S., and M. Braun (1999): "Rentabilidad de la Educacion Formal en Chile," *Cuadernos de Economía*, 36, pp. 685-724.

Beyer, H. (1997): "Distribución del Ingreso: Antecedentes para la Discusión," *Estudios Públicos*, 65 (Verano), pp. 1-54.

Boskin, M. J. (1974): "A Conditional Logit Model of Occupational Choice," *Journal of Political Economy*, 82:2, pp. 389-398.

Brown, C., J. A. Pagán, and E. Rodríguez-Oreggia (1999): "Occupational Attainment and Gender Earnings Differentials in Mexico," *Industrial and Labor Relations Review*, 53(1), pp. 123-135.

Chile Social and Economic Indicators 1960-2000 (2002): *Banco Central de Chile*, Santiago, Chile.

Contreras, D. (2002): "Explaining Wage Inequality in Chile: Does Education Really Matter?" *Journal of Income Distribution*, 11:1, pp. 53-76.

Contreras, D. (2003): "Poverty and Inequality in a Rapid Growth Economy: Chile 1990-1996," *Journal of Development Studies*, 39:3, pp. 181-200.

Cowan, K., A. Micco, A. Mizala, C. Pagés, and P. Romaguera (2005): *Un Diagnóstico del Desempleo en Chile*, Centro de Microdatos, Universidad de Chile.

De Beyer, J., and J. B: Knight (1989): "The Role of Occupation in the Determination of Wages," *Journal of Development Studies*, 41, pp. 595-618.

Ffrench-Davis, R. (2005): *Economic Reforms in Chile, From Dictatorship to Democracy.* Michigan: The University of Michigan Press.

Fields, G., and G. Yoo (2000): "Falling Labor Income Inequality in Korea's Economic Growth: Patterns and Underlying Causes," *Review of Income and Wealth*, 46:2, pp. 139-160.

Gill, I. S., and C. Montenegro (2002): "Responding to Earnings Differentials in Chile," Chapter 6 in *Crafting Labor Policy: Techniques and Lessons from Latin America*, I. S. Gill, C. E. Montenegro, and D. Dömeland (eds.), World Bank and Oxford University Press.

Gindling, T. H. (1991): "Labor Market Segmentation and the Determination of Wage in the Public, Private-Formal, and Informal Sectors in San José, Costa Rica," *Economic Development and Cultural Change*, 39:3, pp. 586-605.

Larrañaga, O. (2001): "Distribucin de Ingresos en Chile: 1958-2001," Department of Economics, Universidad de Chile Working Paper 178.

Lee, L. (1983): "Generalized Econometric Models with Selectivity," *Econometrica*, 51:2, pp. 507-512.

Le, A. T. (1999): "Empirical Studies of Self-employment," *Journal of Economic Surveys*, 13:4, pp. 381-416.

Meller, P., R. O'Ryan, and A. Salomino (1996): "Growth, Equity, and the Environment en Chile: Issues and Evidence," *World Development*, 24:2, pp. 255-272.

Miller, P., and P. Volker (1985): "On the Determination of Occupational Attainment and Mobility," *The Journal of Human Resources*, 20:2, pp. 197-213.

Mizala, A., and P. Romaguera (1996): "Flexibilidad del Mercado del Trabajo: el Impacto del Ajuste y los Requisitos del Crecimiento Económico," Colección Estudios CIEPLAN 43.

Palma, A. (2005): "Size, Significance, and Sources of Recent Income Inequality Changes in Chile," *Working Paper*, Department of Economics, Göteborg University.

Polachek, S. (1979): "Occupational Self-Selection: A Human Capital Approach to Sex Differences in Occupational Structure," *Review of Economics and Statistics*, 63:1, pp. 60-69.

Riveros, L. A. (1990): "The Economic Return to Schooling in Chile: An Analysis of its Long-term Fluctuations," *Economics of Education Review*, 9:2, pp. 111-121.

Saavedra, J., and Al. Chong (1999): "Structural Reforms, Institutions and Earnings: Evidence from Formal and Informal Sector in Urban Peru," *The Journal of Development Studies*, 35:4, pp. 95-116.

Schmidt, P., and R. P. Strauss (1975): "The Prediction of Occupations Using Multiple Logit Models," *International Economic Review*, 16:2, pp. 471-486.

Tiefenthaler, J. (1994): "A Multisector Model of Female Labor Force Participation: Empirical Evidence from Cebu Island, Philippines," *Economic Development and Cultural Change*, 42:4, pp. 719-742.

Uthoff, A. (1986): "Changes in Earnings Inequality and Labor Market Segmentation: Metropolitan Santiago, 1969-1978," *Journal of Development Studies*, 22:2, pp. 300-326.

Yamada, G. (1996): "Urban Informal Employment and Self-employment in Developing Countries: Theory and Evidence," *Economic Development and Cultural Change*, 44:2, pp. 291-314.

Appendix

Figure 1

Occupational Categories

$Non\ \ participate(s = 1)$ $\quad$ $Labor\ \ Market\ \ Choice$

$Self - employment$	$Informal\ \ employment$	$Formal\ \ employment$
$Professionals(s = 2)$		$Professionals(s = 9)$
$Service(s = 3)$	$Service(s = 6)$	$Service(s = 10)$
$Skilled\ \ blue - collars(s = 4)$	$Skilled\ \ blue - collars(s = 7)$	$Skilled\ \ blue - collars(s = 11)$
$Unskilled(s = 5)$	$Unskilled(s = 8)$	$Unskilled(s = 12)$

Table A1

Definitions of the Occupational Groups.

Occupation	*Definition*
Self-Employed	Own-account workers and employers.
Informal workers	Wage and salary workers without written contract.
Formal workers	Wage and salary workers with written contract.
Professionals	Public administrators, managers, professionals, and technicians.
Service	Clericals, sales workers, and other service workers.
Skilled blue-collars	Qualified agricultural workers, officials, operators, and kindred workers.
Unskilled	Unskilled workmen, laborer, and unskilled service workers.

Table A2

Definitions of Variables and Summary Statistics.

Variables	*Definition*	1992	2000
Earnings	Log-monthly earnings of principal occupation.	11.79	12.07
Education	Years of formal education.	9.27	10.34
Ed1	Dummy = 1 if Education $\leq$ 8. otherwise 0.	0.43	0.34
Ed2	Dummy = 1 if Education 9-12. otherwise 0.	0.39	0.41
Ed3	Dummy = 1 if Education 13-16. otherwise 0.	0.12	0.17
Ed4	Dummy = 1 if Education $\geq$17. otherwise 0.	0.06	0.09
Age	Age.	37.94	39.18
Age^2	Age squared and divided by 10.	158.69	168.65
Male	Dummy = 1 if male. otherwise 0.	0.53	0.52
Head	Dummy = 1 if head. otherwise 0.	0.39	0.39
Kids	Dummy = 1 if children ¡ 10 in the household. otherwise 0.	0.58	0.55
Provincial-resident	Dummy = 1 if residents outside Santiago. otherwise 0.	0.59	0.59
Unemployed	Dummy = 1 if anyone unemployed in the household. otherwise 0.	0.07	0.11

Table A3

Percentage Working Part-time, 1992 and 2000

Sector	*Occupation*	*1992* (1)	*2000* (2)	*Ratio* (2)/(1)
Self-Employment	Professionals	15.79	19.34	1.22
	Service	19.67	31.94	1.62
	Skilled blue-collar	19.63	26.34	1.34
	Unskilled	29.29	39.59	1.35
Informal Employment	Professionals	13.30	15.93	1.20
	Service	13.21	29.22	2.21
	Skilled blue-collar	8.78	14.00	1.59
	Unskilled	15.63	33.13	2.12
Formal Employment	Professionals	13.30	15.93	1.20
	Service	3.96	8.30	2.10
	Skilled blue-collar	2.17	5.55	2.56
	Unskilled	6.69	7.33	1.10
Self-Employment		20.36	26.22	1.29
Informal Employment		13.23	26.08	1.97
Formal Employment		6.07	9.11	1.50
Total		**6.89**	**10.26**	**1.49**

Notes: Columns (1) and (2) represent the percentage of part-time workers within respective occupation. Columns (2) and (4) represent the percentage of part-time work within each sector.

Table A4

Distribution by Occupational Group, 1992 and 2000

Sector	*Occupation*	*Males*		*Females*	
		1992	*2000*	*1992*	*2000*
		(1)	(2)	(3)	(4)
Inactive		3.69	14.26	60.71	56.56
Active		96.31	85.74	39.29	43.44
Total		**100**	**100**	**100**	**100**
			Among Employed Workers		
Self-Employment	Professionals	6.01	8.38	5.84	7.81
	Service	2.95	1.87	7.65	5.59
	Skilled blue-collar	14.03	13.59	4.62	4.22
	Unskilled	4.68	2.70	3.53	2.84
Informal Employment	Service	1.05	1.63	3.21	6.36
	Skilled blue-collar	3.48	6.82	0.98	1.43
	Unskilled	4.57	4.96	5.92	10.31
Formal Employment	Professionals	11.63	14.93	18.63	21.97
	Service	10.25	11.12	27.14	26.38
	Skilled blue-collar	25.46	22.25	7.14	3.88
	Unskilled	15.90	11.76	15.32	9.21
Self-Employment		27.67	26.54	21.64	20.46
Informal Employment		9.09	13.40	10.11	18.09
Formal Employment		63.24	60.06	68.24	61.44
Total		**100**	**100**	**100**	**100**

Table A5

Earnings Inequality by Occupational Group, 1992 and 2000

Sector	*Occupation*	*Males*		*Females*	
		1992	*2000*	*1992*	*2000*
		(1)	(2)	(3)	(4)
Self-Employment	Professionals	1.29	1.27	1.16	1.04
	Service	0.70	0.75	0.80	0.84
	Skilled blue-collar	0.68	0.64	1.01	0.93
	Unskilled	0.64	0.49	0.80	0.77
Informal Employment	Service	0.32	0.51	0.42	0.54
	Skilled blue-collar	0.36	0.36	0.31	0.42
	Unskilled	0.31	0.36	0.37	0.42
Formal Employment	Professionals	0.72	0.69	0.53	0.47
	Service	0.37	0.29	0.33	0.39
	Skilled blue-collar	0.33	0.27	0.24	0.21
	Unskilled	0.25	0.15	0.23	0.18
Self-Employment		1.06	1.19	1.22	1.19
Informal Employment		0.35	0.41	0.41	0.48
Formal Employment		0.59	0.58	0.50	0.55
Total		**0.77**	**0.82**	**0.71**	**0.78**

Table A6

Average Earnings by Occupational Group, 1992 and 2000

Sector	*Occupation*	*Males*		*Females*	
		1992	*2000*	*1992*	*2000*
		(1)	(2)	(3)	(4)
Self-Employment	Professionals	994	1200	573	502
	Service	326	314	221	202
	Skilled blue-collar	251	257	166	165
	Unskilled	164	162	103	108
Informal Employment	Service	109	140	86	106
	Skilled blue-collar	104	119	84	86
	Unskilled	75	83	58	72
Formal Employment	Professionals	463	621	246	368
	Service	166	213	129	164
	Skilled blue-collar	144	184	101	122
	Unskilled	93	124	69	105
Self-Employment		405	570	284	297
Informal Employment		90	108	69	85
Formal Employment		193	287	144	226
Total		**242**	**334**	**167**	**216**

Notes: Values represent thousands of Chilean pesos of 2000.

Table A7

Total Decomposition of Changes in Earnings-Inequality, by Gender and Sector, 1992-2000

	Within-group		*Between-group*		***Total***
	Composition-effect	*Change in variance*	*Composition-effect*	*Change in variance*	***contribution***
			Males		
Self-Employment	0.74	-1.01	3.50	-2.06	**1.17**
Informal Employment	1.49	0.62	1.21	3.18	**6.50**
Formal Employment	0.61	-3.72	0.79	-0.15	**-2.48**
Total contribution	**2.84**	**-4.11**	**5.49**	**0.97**	**5.19**
			Decompocition Within-sectors		
Self-Employment	6.22	-3.81	5.49	4.60	**12.49**
Informal Employment	0.58	4.63	-0.01	1.06	**6.27**
Formal Employment	3.07	-6.20	3.61	-0.99	**-0.52**
			Females		
Self-Employment	-0.31	-1.21	1.87	-3.94	**-3.59**
Informal Employment	3.06	1.49	3.37	2.51	**10.43**
Formal Employment	-0.68	-0.35	-0.97	3.38	**1.37**
Total contribution	**2.07**	**-0.08**	**4.27**	**1.95**	**8.22**
			Decompocition Within-sectors		
Self-Employment	3.94	-5.94	-0.47	4.35	**1.89**
Informal Employment	0.26	8.21	-0.18	-1.85	**6.45**
Formal Employment	2.79	-0.57	-2.65	4.59	**4.17**

Notes: All numbers have been multiplied by 100.

Table A8

Relative Factor Inequality Weight, 1992

		Education	*Age*	*Male*	*Provincial resident*	*lnHour*	λ_{so}	**Total**
Self-Employment	Professionals	0.196	0.025	0.016	0.004	0.035	0.044	**0.320**
	Service	0.060	0.007	0.042	0.005	0.075	0.002	**0.190**
	Skilled blue-collar	0.082	0.015	0.060	0.021	0.076	-0.006	**0.247**
	Unskilled	0.034	0.008	0.054	0.026	0.112	0.003	**0.237**
Informal Employment	Service	0.051	0.003	0.034	0.003	0.119	0.000	**0.211**
	Skilled blue-collar	0.028	0.005	0.014	0.035	0.048	-0.002	**0.128**
	Unskilled	0.004	0.000	0.060	0.037	0.107	-0.001	**0.209**
Formal Employment	Professionals	0.121	0.060	0.085	0.048	0.053	0.002	**0.370**
	Service	0.116	0.053	0.033	0.018	0.002	-0.003	**0.220**
	Skilled blue-collar	0.073	0.043	0.034	0.007	0.011	-0.008	**0.161**
	Unskilled	0.029	0.000	0.082	0.022	0.037	-0.004	**0.166**

Source: Author's calculations from CASEN.

Table A9

Relative Factor Inequality Weight, 2000

		Education	*Age*	*Male*	*Provincial resident*	*lnHour*	λ_{so}	**Total**
Self-Employment	Professionals	0.191	0.013	0.032	0.016	0.026	0.171	**0.450**
	Service	0.092	0.010	0.029	0.001	0.199	0.004	**0.336**
	Skilled blue-collar	0.106	0.016	0.045	0.024	0.153	0.003	**0.346**
	Unskilled	0.028	0.007	0.051	0.007	0.301	-0.003	**0.391**
Informal Employment	Service	0.068	0.008	0.005	0.032	0.226	0.012	**0.352**
	Skilled blue-collar	0.041	0.007	0.038	0.042	0.210	-0.002	**0.336**
	Unskilled	0.018	0.002	0.009	0.044	0.345	0.002	**0.420**
Formal Employment	Professionals	0.124	0.052	0.047	0.030	0.068	0.065	**0.386**
	Service	0.113	0.029	0.032	0.014	0.143	0.000	**0.331**
	Skilled blue-collar	0.140	0.033	0.049	0.006	0.019	-0.012	**0.236**
	Unskilled	0.048	0.004	0.051	0.055	0.064	0.004	**0.227**

Source: Author's calculations from CASEN.

Price and Composition Effects on the Rise and Fall of Wages Inequality in Santiago

ALEXIS PALMA †‡

Abstract

This paper studies the development of wage inequality in Santiago's labor market during the period 1970-2003 using a Oaxaca-Blinder type decomposition based on quintile regressions suggested by Machado and Mata (2004). I calculate counterfactual distributions to analyze the price and composition effects on inequality changes in the mentioned period.

Log-wage inequality increased very significantly between 1974 and 1987 before it declined greatly between 1987 and 1992. In the following years inequality was more stable than in previous decades but is characterized by an inverted-U pattern between 1992 and 2003. This pattern also characterizes the inequality in the upper part of the wage distribution.

I found significant changes in inequality making consecutive comparisons of 1974, 1987, 1992, 2000, and 2003 which represent extreme values in the distribution of wages during the analyzed period.

The great deterioration in inequality between 1974 and 1987 was driven by a significant prices and composition change, both of which had an inequality increasing effect. The following inequality decline, between 1987 and 1992, is mainly explained by the inequality decreasing effect of prices, which explains the totality of the first and fifth deciles' change between these years.

†Department of Economics, Göteborg University, Vasagatan 1, SE-405 30 Göteborg, Sweden. Phone: +46 (0)31 786 13 67. E-mail: alexis.palma@economics.gu.se

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1 Introduction

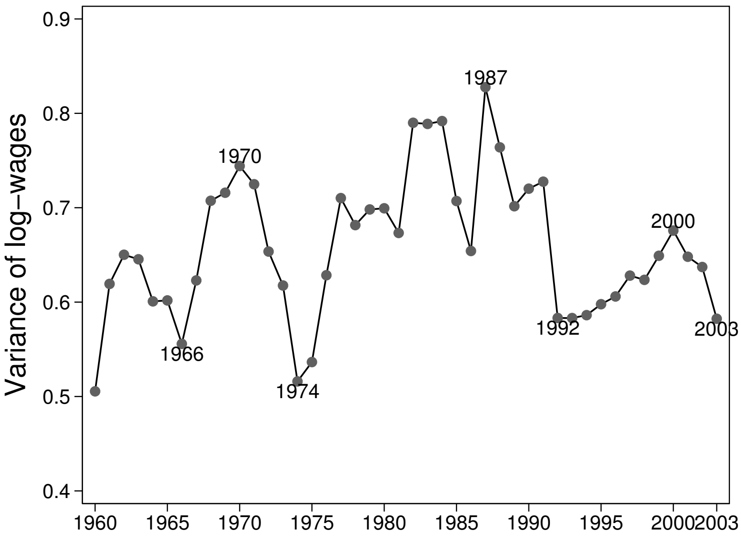
It is a well established stylized fact that the 1980s was a decade characterized by great changes in wage inequality in almost all industrial countries (Gottschalk and Smeeding, 1997). Substantial research has been done and several explanations, such as a skill-biased technological change and increased international trade, have been given to this increased inequality.

Rising inequality patterns were also observed in many developing countries (Psacharopoulos *et al.*, 1995; Krongkaew, 1994; Zin, 2005), but the explanations for this stylized fact have not received much attention in the international literature. This paper will focus on Chile, a country where inequality, defined as the variance of log-wages of male workers, fluctuated greatly during the last four decades. Figure 1 below shows that after have declined for several years in the early 1970s, inequality increased considerably between 1975 and 1977. Beginning with a major deterioration in the 1975-77 period, between 1977 and 1992 the variance of log-wages fluctuated between 0.68 and 0.82 with top values found in 1982-84 and 1987. The period ended with a 20% decline in the degree of inequality. In the following years of the research period inequality was lower, and fluctuated less, than in the previous decade, but followed an inverted-U pattern between 1992 and 2003 with a peak in year 2000. However, inequality behaved quite different in the different parts of the log-wage distribution.

In Figure 2, inequality is defined as $Q(0.90) - Q(0.50)$ and $Q(0.50) - Q(0.10)$, representing the difference between the ninth decile and the median of the wage distribution, and the median and first decile, respectively. I set 1960=1.0 to make the two series comparable. This figure shows that the inequality in the upper part of the wage distribution followed in general a similar pattern than the variance of

Figure 1

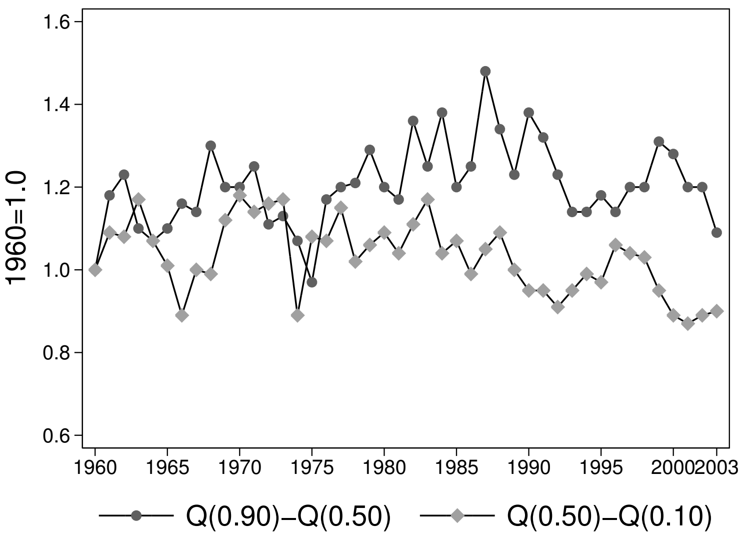
Variance of Log-Wage, 1960-2003



Source: Author's calculation from Employment and Unemployment Survey of Universidad de Chile for Greater Santiago.

Figure 2

Log-Wage Inequality in the Different Parts of the Distribution, 1960-2003



Source: Author's calculation from Employment and Unemployment Survey for Santiago.

log-wages, increasing from approximately 1.0 in the mid 1970s to 1.4 in the late 1980s. This deterioration in the inequality in the upper part of the distribution was followed by a decline between 1987 and 1992. In the lower part of the wage distribution, on the other hand, inequality followed a different pattern fluctuating less than in the upper part of the distribution and being almost constant over the research period. There is also evidence in the literature (Palma, 2005) that an inequality increase, between the early 1970s and the late 1980s, was also observed for white-collars, own-account workers, employers, and female workers; but for males the increase (18%) was one of the largest.

The source of the fluctuations in the inequality of wages in Chile has been studied in different works (Contreras, 2002; Beyer *et al.*, 1999). Their results suggest that education and trade liberalization had an inequality increasing effect in Chile's labor market. Moreover, other research (Meller *et al.*, 1996; Larrañaga, 2001) indicate that a higher rate of unemployment and a higher rate of participation of women in the labor force tend to deteriorate the inequality of household income.

Despite these insights, several pieces of information are still missing in the literature. Whether the sources that drive inequality in the right tail of the wage distribution are different from those in the left tail has not been carefully analyzed yet. Nor has the effect of the change in the population structure of employment observed in last decades on inequality been carefully scrutinized. These two issues are addressed in this paper. Using a quantile regression approach developed recently by Machado and Mata (2004), I generate counterfactual distributions of wages, a method that has not so far has not been applied on Chilean data. The value of this method is that allows us to asses the effect of changes in the parameters of the wage equation and of covariates across the entire distribution of wages. I focus on five years that represent extreme cases of inequality displayed in the figures above, namely 1974, 1987, 1992, 2000, and 2003.

My results indicate that significant changes in inequality occurred in 1987, 1992, 2000, and 2003. Inequality increased between 1974 and 1987 because of the inequality increasing effects of prices and composition change, both of which were highly significant according to the Bootstrap analysis. Underlying the inequality decline between 1987 and 1992, is the inequality decreasing effect of prices, which explains the totality of the first and fifth deciles' change.

The rest of the paper is structured as follows. Section 2 presents a background discussion. Section 3 presents the methodology and the inequality decomposition. In section 4 I present a description of the data used in the paper, followed by the result. In the last section, I present a summary and my conclusions.

2 Background

The Chilean economy went through several dramatic political and economic changes during the decades covered in this study. From being an economy characterized by a upsurge in the participation of the state, several price regulations, import substitution, and historically low rates of unemployment in the early 1970s, it changed towards a liberalized, outward-oriented economy with high unemployment rates during the second half of the 1970s and the 1980s (Table 1). The 1990s represent a period of unprecedented high rate of growths in GDP (Gross Domestic Product) thanks to a strong expansion of exports and investments. In the face of the substantial expansion of output during the 1990s, the period 1991-2000 is characterized by a 8% rate of unemployment, five percentage points below the unemployment rate of the previous period.

Especially dramatic during the research period were the changes in labor market policies. After the increased participation of labor unions and a growing job security in the late 1960s and the early 1970s, the military administration banned unions, suspended collective bargaining, and induced a *de facto* reduction of job security in the period 1974-79 (Cortázar, 1997; Edwards and Edwards, 2000). Along with these policies, wage readjustments were put under the strict control of the government authorities. The results were a proliferation of unemployment and a serious deterioration of the minimum wage. But since one of the main focuses of the economic policy were put on the fight of inflation, average wage turned positive.

Table 1

Some Labor Market Indicators, 1970-2000

(All numbers represent %)

	Union density	Collective bargaining	Real average wage growth (per year)	Real minimum wage growth (per year)	Unemployment rate
	(1)	(2)	(3)	(4)	(5)
1970-1973[a]	34.7	26.1	-5.1	35.3	4.7
1974-1979	-	-	0.9	-9.2	13.7
1980-1990	8.0	13.4	1.3	-0.7	14.7
1991-2000[b]	15.0	16.7	3.6	5.7	8.5

Sources: (1) Edwards and Edwards (2000); (2)-(4) Cortázar (1997) and Ffrench-Davis (2005); (5) Chile Social and Economic Indicators 1960-2000.

Notes: (1) Percentage of wage earners who were union members; (2) percentage of wage earners covered by collective agreements; [a] refers only to 1970 for union density; [b] refers only to the period 1991-1993 for collective bargaining.

The new labor code introduced in 1980 allowed unions but under some restrictions and collective bargaining was only permitted at the firm-level. Moreover, a series of restrictions that controlled the possibility to work in many occupations, firms' possibility to subcontract, and the control of the authorities on the wage adjustments were removed. And despite that almost two decades had pass, the average real wage was 25% lower in 1990 compared with 1971.

The few changes in the labor code during the 1990s increased job security and strengthened the bargaining position of unions. Union density and collective bargaining recovered but never to pre-Pinochet levels. This period is also characterized by a rapid expansion of the real average and the real minimum wage.

In the Chilean literature, the increased return to university education that followed the liberalization of the labor market and the opening of the economy have been suggested to be one of the main explanations to the high level of inequality in the 1980s. For instance, Contreras (2002) suggests that, on average, education explained 38% of the wage inequality of full-time male workers; the greatest part of this share arises from the return to university education which increased from on average 0.16 in the early 1970s to 0.26 in the late 1980s. Another relevant work is Montenegro (1998), who found that the rate of return to education increased in the upper part of the wage distribution and decreased in the lower part during the period 1975-87. Scrutinizing narrower groups of the labor market, Montenegro found a strongly increasing pattern of the rate of return to education for less experienced and white-collar workers. Other groups such as experienced, public, and private sector workers, followed the same pattern, though to a lesser degree. Montenegro also shows that inequality (defined as $S_{(90,10)} = \text{antilog}(Q(0.90) - Q(0.10)) - 1$) increased from the mid 1970s until 1987 for the total sample, which was also the case for experienced, private, and white-collar workers as well. Using instead $S_{(75,25)}$ as the definition of inequality, he found a much more stable inequality, and none of the previous mentioned groups reported major changes over the research period. Montenegro's results, together with

Figure 2, suggest that understanding the behavior of the inequality in the upper part of the distribution is crucial to understand how overall inequality behaves over time.

3 Methodology

The method I use in this paper is an extension of the well-known Oaxaca decomposition used to analyze the effect of differences in the mean of covariates (the composition effect) and of the parameters of the wage equation (the price effect) on the mean of two different wage distributions. However, in this paper this method is used to analyze the entire wage distribution. Thus, I assume that inequality changes can arise because of changes in the remuneration of individual characteristics, such as education and experience, or because the distribution of characteristics changes. Using the entire distribution, one is able to perform an inequality decomposition with any inequality indicator and explain inequality changes in different parts of the distribution. This method has been used in recent years by Autor *et al.* (2005) and Machado and Mata (2004), but so far has not been applied to Chilean data.

The methodology is the following: let $Q_\theta^t(y|z)$ for $\theta \in (0,1)$ represents the $\theta : th$ quintile of the distribution of wages, y, given the covariates z in year t. The model of the conditional quantiles is then

$$Q_\theta^t(y|z) = z_t'\beta_t(\theta) \tag{1}$$

where $\beta_t(\theta)$ is a vector of coefficients of the θ:th quintile regression. This regression model estimates the θ percentile of the dependent variable, conditional on the values of the explanatory variables, in the same way that OLS estimates the mean of the dependent variable.

This inequality decomposition is based on the idea of the probability integral transformation theorem which suggest that if $\theta_1, \theta_2, ..., \theta_m$ are randomly selected numbers from the unit interval $[0,1]$, the corresponding m estimates of the conditional quantile of wages at z_t, $\{z_t' b_t(\theta_j)\}_{j=1}^m$, form a random sample from the (estimated) conditional distribution of wages given by z_t. As I am interested in the unconditional distribution of male wages, I follow Machado and Mata who suggest to 'integrate z out' by selecting a random sample of the covariates from an appropriate distribution. The procedure is the following:

1. Draw m numbers at random from (0,1) and denote these numbers $u_1, u_2, ..., u_m$.

2. Using the data set Z_t, estimate the quantile regressions for period t

$$Q_{u_j}^t(y|z) = z_t' \beta_t(u_j), \tag{2}$$

this generates m vectors of estimated quintile regression parameters $b_t(u_j)$.

3. Make m draws at random with replacement from the data set of year t, Z_t. Denote the j vector of that data $z_t^*(j)$

4. Then:

$$y_t^*(j) = z_t^*(j)' b_t(u_j), \qquad \text{for} \qquad j = 1, ..., m \tag{3}$$

are m elements of the desired distribution. The distribution of this variable is denoted $f^*(y_t)$.

To simulate the counterfactual distribution of wages when individual characteristics are paid as in year t, but the covariates are distributed as in year s, I use Z_s in step 3 and denote this distribution $f^*(B_t, Z_s)$. To generate the counterfactual distribution when only one covariate is distributed as in year s, I proceed in the following way. First, consider the occupational groups denoted O_1=blue-collar workers, O_2=white-collar workers, O_3= professionals.

a. Generate m elements of the wage distribution $\{y_t^*(j)\}_{j=1}^m$ in year t according to the method described above.

b. Select the sample of $\{y_t^*(j)\}_{j=1}^m$ of the individuals that are blue-collar workers.

c. Generate a random sample of size $m * p_s^{O_1}$, where $p_s^{O_1}$ is the relative frequency of blue-collar workers in year s, with replacement from the sample drawn in step b.

d. Repeat steps b and c for the other two occupational groups.

In this way, one re-weights the distribution of wages of year t using the weights of the occupational groups of year s to obtain the counterfactual distribution. This distribution is denoted $f^*(B_t; O_s)$.

The inequality decomposition is then the following:

$$I(f(y_t)) - I(f(y_s)) = I(f^*(B_t; Z_s)) - I(f^*(y_s)) + I(f^*(y_t)) - I(f^*(B_t; Z_s)) + Residuals,$$

$$I(f(y_t)) - I(f(y_s)) = \text{Price effect+Composition effect+Residuals.}$$

$I(f(y_t))$ represents the observed inequality indicator based on the distribution $f(y_t)$ of the wages of year t. On the left-hand side, one has the observed change in inequality between the two periods. On the right-hand side, the first term is price effect on inequality change. This is the difference between the inequality of wages when the distribution of individual characteristics is the same in both periods, but all the parameters of the wage equation are changed. The second term is the effect of change in the composition of the workers, and the third is a residual term. If one is interested only in the contribution of an individual covariate, O, on the inequality change, one calculates:

$$\begin{array}{l}\text{Contribution of a change in a single} \\ \text{covariate to the inequality change}\end{array} = I(f^*(y_t)) - I(f^*(B_t; O_s)).$$

I calculate 95% confidence intervals for the different components of the decomposition using a Bootstrap technique and 100 replications.

4 Data Description

Employment and Unemployment Survey of Universidad de Chile for Greater Santiago (EUSS) is a household survey collected in the Metropolitan region, where Santiago is located and roughly 40% of the Chilean population lives. EUSS has been carried out every year since 1957 by Universidad de Chile. As its name hints, the main purposes of the survey is to obtain information about some labor market variables for the Metropolitan region, but also information about incomes is collected. The selected households were visited four times, namely in March, June, September, and December; but only in June EUSS collects information about incomes. The survey was initiated with 3.500 households in the late 1950s but because of different reasons it declined to 3.400 in 1974, and to 3.060 in 1983. Since then the number of households surveyed has remained constant.

The selection process is of multistages. Great Santiago's different municipalities are grouped into eight strata. Within each strata blocks of dwellings, or groups of blocks of dwellings, form the second level strata which are selected with probability related to their size. The final level strata are groups of approximately ten private households. Moreover, the selection process has been adjusted year by year to reflect the changing size and geographical structure of Great Santiago.

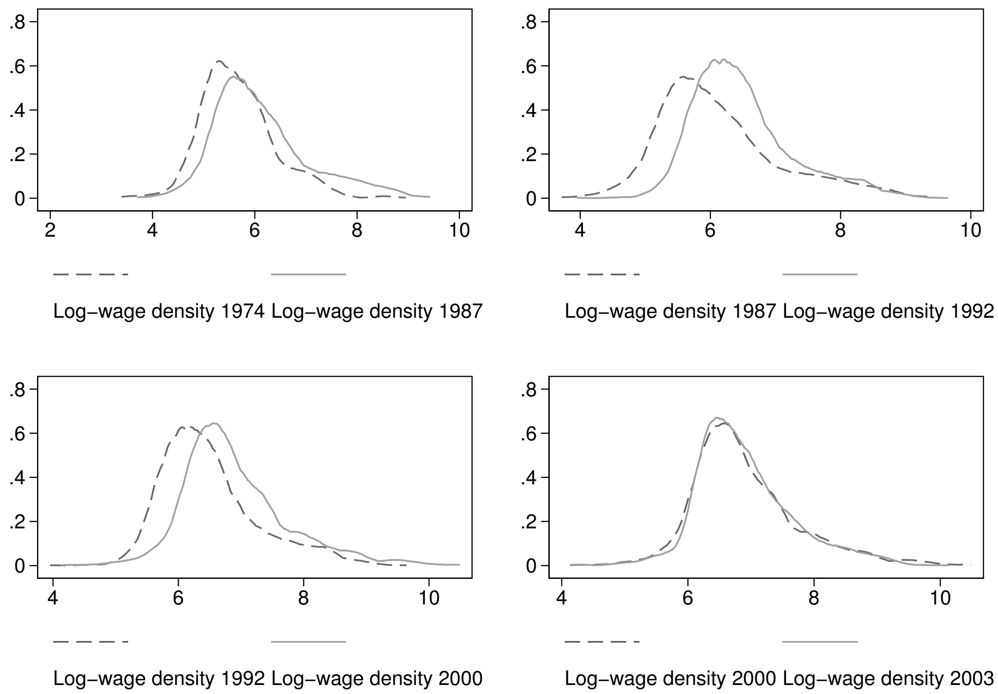
Changes in the questionnaire were introduced in 1980 when a new question about the willingness to work was introduced in order to obtain a more accurate measure of unemployment. Additionally, from 1998 the question about the level of education of the individual was changed to obtain more detailed information about the type of education. Beside of that, the questions in the survey have not been changed over the years, therefore EUSS is especially appropriate for long-run analysis. This paper considers only male wage and salary workers, in order to avoid the issue of labor force participation. The Self-employed are excluded since one is not able to distinguish between return to labor and return to capital from the total income of these individuals. Left is a sample of approximately 1700 individual for each of the rounds of EUSS used here. The variable I analyze is labor income. These incomes were adjusted by Universidad de Chile for missing values according to the Hot Deck method.

Figure 3 shows the estimated Kernel density of the log-wages for the five waves of EUSS I use. The figures show clearly that the distribution of wages shifted rightwards as real wages increased. The shift of the distribution represents an increase of real wages (in natural units) of 46% between 1974 and 1987. It is worth noting that in the following periods, that covers only five and eight years, and which represents the period with high GDP-growth rates, the mean of the distribution increased by 37% and by 41%, respectively.

In the last period the mean wage was virtually constant, reflecting the deceleration of the Chilean economy that one observes at the end of the 1990s. Behind average values, the wage growth rate was considerably higher in the right tail of the distribution comparing 1987 with 1974, when the ninth decile increased by 80%, compared with 30% and 20% for the fifth and first deciles, respectively. These figures indicate that the relatively low growth rate of wages during these years did not characterize the entire distribution only the lower parts of it. This disproportionately high growth rate in the right tail of the distribution coincides with the period when Chile became more open to international trade, which allowed the country not only to increase but also to diversify its exports.

Figure 3

Kernel Log-wage Densities for Years 1974, 1987, 1992, 2000, and 2003



Source: Author's calculation from EUSS.

Notes: Wages are expressed in constant prices of December of 1998.

Several studies have suggested that in the Chilean case this process was closely associated with a growing demand for skilled workers, generating also a concomitant increase in the skill premium (Robbins, 1994). The result was a marked expansion in wage inequality between 1974 and 1987 (1987 is actually the year with the highest inequality of the research period, as Table 2 shows).

In 1992, inequality declined considerably, following an inverted-U pattern in the remaining years of the research period. Between 1987 and 1992, not only was the growth rate of the first and fifth deciles larger than in the previous period (50% for the first decile and 40% for the fifth), it was considerable smaller for the ninth decile (only 8%).

Table 2

Male Wage Inequality Indicators for different Years

	1974	1987	1992	2000	2003
	(1)	(2)	(3)	(4)	(5)
Variance of log-wages	0.528	0.886	0.602	0.705	0.581
(SE)	(0.025)	(0.034)	(0.022)	(0.029)	(0.024)
$Q(0.10)$ of log-wages	4.900	5.137	5.661	6.080	6.122
(SE)	(0.028)	(0.032)	(0.015)	(0.021)	(0.018)
$Q(0.50)$ of log-wages	5.593	5.914	6.355	6.746	6.772
(SE)	(0.019)	(0.016)	(0.034)	(0.022)	(0.029)
$Q(0.90)$ of log-wages	6.691	7.560	7.646	8.047	7.911
(SE)	(0.046)	(0.071)	(0.057)	(0.052)	(0.050)
Gini coefficient of wages	0.441	0.567	0.477	0.528	0.459
(SE)	(0.013)	(0.010)	(0.009)	(0.014)	(0.009)
Theil index of wages	0.396	0.663	0.441	0.605	0.419
(SE)	(0.032)	(0.029)	(0.021)	(0.042)	(0.025)

Sources: Author's calculations from EUSS.
Notes: The numbers are calculated using only male wage and salary workers from EUSS surveys. Standard error, (SE), in parenthesis are calculated using 100 Bootstrap replications. Wages are expressed in constant prices of December 1998.

This was the period when the Chilean economy starts its most prolonged period of GDP-growth, initially supported by the expansion of exports and the improvement in the terms of trade, but later on also by enlarged investments. At the same time minimum wage increased sharply between 1989 and 1993, which might be one important factor to explain the increased wages in the left tail of the distribution.

The rest of the 1990s is the only period when wages increase at the same rate (40%) over the whole distribution.

I also present two other inequality indicators; namely the Gini coefficient and the Theil inequality index, to see if they tell the same story. In relative terms, the changes of the Gini coefficient are smaller than of the Log-variance. The Theil inequality index, on the other hand, reports similar changes to the ones reported by the Log-variance with exception of 2000-03, when it reports larger relative changes than those given by the Log-variance and the Gini coefficient.

The significance of income inequality changes has received almost no attention in previous Chilean studies, but starts to be a standard procedure in the analysis of income inequality changes in the international literature (see for example Athanasopoulos and Vahid, 2003; and Moran, 2006). Because of this, I apply the Bootstrap technique suggested by Mills and Zandvakilly (1997) to study the difference between the inequality of period t, $(I(t))$, and of period s, $(I(s))$. If the Bootstrap confidence interval of the variable $\Delta I = I(t) - I(s)$ does not include zero, then one can draw the conclusion that the inequality change is significant. Applying this technique I found that a statistically significant Log-variance change occurred for all consecutive periods I study (see Tables A2-A4). These results are corroborated using the Gini coefficient as inequality indicator. Not surprisingly, the inequality difference comparing 1974 with 1992 and 2003 were not significant since the level of earnings inequality was very similar in these three years.

Information on how the population structure of the individuals I study has changed, and which is the information that I used to calculate the counterfactual distributions, is found in Table 3.

Table 3

Descriptive Statistics of Covariates, 1974-2003

	1974 (1)	1987 (2)	1992 (3)	2000 (4)	2003 (5)
Log-wage					
Mean	5.66	6.12	6.49	6.90	6.89
Standard deviation	0.72	0.94	0.77	0.84	0.76
Age structure (%)					
≤24	22.6	20.9	18.5	15.2	12.7
25-34	33.1	32.7	36.1	29.0	29.8
35-44	23.1	23.9	20.4	27.6	27.7
45-54	13.6	14.3	15.2	16.4	18.6
55≤	7.5	7.9	9.7	11.5	11.0
Education (%)					
Primary Educated	56.1	31.9	26.4	22.1	19.9
Secondary Educated	34.1	50.0	58.5	59.8	61.4
University Educated	9.7	17.7	14.8	17.9	18.6
Occupation (%)					
Blue-collar	54.2	54.0	55.0	56.2	56.2
White-collar	33.0	29.6	28.6	25.8	26.8
Professionals	12.7	16.2	16.2	17.9	16.9
Sector employment (%)					
Manufacturing	39.4	29.3	33.5	26.0	23.8
Construction	14.5	11.7	13.9	13.5	13.1
Commerce	10.6	15.5	13.9	16.9	17.1
Financial Services	7.7	13.9	11.2	14.0	14.9
Personal Services	5.7	6.1	5.7	5.3	5.1
Communal Services	10.2	12.3	10.8	12.0	12.1
Transport	11.6	10.9	10.7	11.9	13.6

Sources: Author's calculations from EUSS.

Notes: The numbers are calculated using only male wage and salaried workers from EUSS surveys. Wages are expressed in constant prices of December 1998.

This table shows that the age structure changed towards older workers (older than 35), at the same time the level of education increased through a doubling of the percentage of male workers with secondary and university education. This was the result of a new education policy introduced in the early 1980s which increased the participation of the private sector in providing these levels of education. This together with the introduction of tuition fees provided additional found for universities to increase the number of place available.

The employment structure across occupations changed slightly towards blue-collar and professional workers, with most of the changes occurring between 1987 and 2000 for the former group and between 1974 and 1987 for the latter group of workers. The percentage of male workers employed in manufacturing declined strongly between 1974 and 1987 and after 1992. This deindustrialization implied that the population share of this sector was 16 percentage points lower in 2003 compared with 1974. Another sector that clearly declined in the period 1974-87 was construction, while service sectors such as commercial, financial, personal, and communal services increased. Exactly the opposite pattern is seen in the period 1987-92, when manufacturing and construction increased and the other sectors declined. In the remaining periods employment in service sectors expanded considerably, accounting for 62% of the total employment in 2003, compared with 52% in 1992.

5 Quintile Regression Results

After having provided information on the methodology and data used, I proceed to present the results of the quantile regressions and of the inequality decomposition. The dependent variable is the monthly earnings from principle occupation divided by the number of hours worked, in log form. The first explanatory variables are age and age squared used as a proxy for in-work experience. The data lack information about the years of schooling but contain information on the level of education of the individuals. I therefore use two dummy variables to capture the effect of secondary and university education on the wage rate. I include two dummies to analyze the effect of white-collar and professional occupations as several studies have highlighted that the trade liberalization has had a disproportionately positive effect on the income of skilled occupations and on the distribution of their earnings. Also dummies to control the effect of occupation in the different economic sectors are included as regressors. The results of the first, fifth, and ninth decile regressions and the OLS estimates are found in Table 4.

Most of the parameters have the expected sign, and the explanatory power of the regression is higher for the ninth decile than for the first. The linear and quadratic terms in age have the expected positive and negative sign, respectively, in all regressions. Nevertheless, the parameters of age squared were not significant for the ninth decile in 1992 and 2000. The curvature of the wage-experience profile in the first decile was very similar over the years, reaching a top at the age of 45. The exception was 1992 when the coefficient of age declined considerable and the wage-experience profile become almost a straight line. This indicate that for the period from 1992, age become a worse proxy for experience. For the nine decile the top increased over the years to be an almost straight line, although no significant, for the whole period after 1992.

The coefficients of the quintile regression reveal also that the secondary education wage-premium increases as we move up in the wage distribution becoming in the ninth decile twice as high as in the first decile. This was the case in 1974, 1987, and in 2000, but when inequality declined in 1992, this difference was clearly reduced. Similar inter-quintile differences were found for the university education wage-premium but predominantly in 1974, 1987 and 2003.

Table 4

Estimates of the Quantile Regressions, (Dependent Variable: Log-Wage)

	1974				1987				1992			
	$b(0.1)$	$b(0.5)$	$b(0.9)$	OLS	$b(0.1)$	$b(0.5)$	$b(0.9)$	OLS	$b(0.1)$	$b(0.5)$	$b(0.9)$	OLS
Constant	3.728***	4.038***	4.322***	3.938***	3.963***	4.376***	4.776***	4.381***	4.780***	5.230***	5.649***	5.246***
	(17.054)	(32.707)	(23.923)	(32.845)	(16.463)	(31.301)	(19.037)	(31.179)	(26.020)	(43.635)	(21.723)	(45.330)
Age	0.062***	0.067***	0.083***	0.073***	0.053***	0.055***	0.066***	0.055***	0.033***	0.032***	0.033**	0.031***
	(5.268)	(10.077)	(8.776)	(11.284)	(4.249)	(7.547)	(5.485)	(7.502)	(3.618)	(5.409)	(2.416)	(5.372)
Age^2	-0.007***	-0.007***	-0.009***	-0.008***	-0.006***	-0.005***	-0.006***	-0.005***	-0.003**	-0.003***	-0.002	-0.002***
	(5.056)	(8.697)	(7.495)	(9.813)	(3.749)	(5.798)	(4.405)	(5.722)	(2.576)	(3.710)	(1.139)	(3.351)
Secondary	0.160**	0.264***	0.347***	0.246***	0.169***	0.237***	0.274***	0.220***	0.258***	0.264***	0.280***	0.269***
	(2.307)	(7.335)	(6.444)	(7.166)	(2.649)	(6.403)	(4.126)	(5.922)	(4.877)	(7.657)	(4.141)	(8.093)
University	0.386***	0.779***	1.021***	0.746***	0.816***	1.257***	1.328***	1.074***	0.714***	1.080***	1.026***	0.966***
	(3.025)	(11.927)	(10.973)	(11.874)	(7.098)	(21.368)	(13.056)	(18.170)	(7.749)	(18.801)	(8.691)	(17.436)
White-collar	0.163**	0.316***	0.420***	0.321***	0.412***	0.521***	0.583***	0.513***	0.223***	0.335***	0.677***	0.396***
	(2.200)	(8.319)	(7.365)	(8.827)	(5.920)	(13.430)	(8.330)	(13.167)	(3.836)	(9.342)	(9.185)	(11.483)
Professional	0.505***	0.628***	0.631***	0.622***	0.804***	0.959***	1.176***	1.047***	0.516***	0.650***	0.970***	0.732***
	(4.154)	(10.067)	(7.318)	(10.397)	(7.209)	(16.636)	(11.898)	(18.096)	(5.809)	(12.142)	(8.729)	(14.201)
Construction	-0.109	-0.036	-0.114*	-0.044	-0.063	-0.070	-0.047	-0.028	-0.146**	0.001	0.004	-0.043
	(1.368)	(0.799)	(1.648)	(1.040)	(0.718)	(1.373)	(0.571)	(0.538)	(2.181)	(0.022)	(0.050)	(1.019)
Commerce	-0.197**	-0.106**	0.025	-0.085*	-0.376***	-0.186***	-0.168**	-0.189***	-0.230***	-0.112**	-0.096	-0.146***
	(2.174)	(2.123)	(0.327)	(1.796)	(4.790)	(4.054)	(2.074)	(4.099)	(3.396)	(2.568)	(1.108)	(3.488)
Financial	-0.069	-0.012	-0.180**	-0.096*	-0.179**	-0.156***	-0.078	-0.094*	0.103	0.080*	-0.031	0.092**
	(0.650)	(0.215)	(2.105)	(1.753)	(2.145)	(3.241)	(1.008)	(1.939)	(1.297)	(1.647)	(0.318)	(1.969)
Personal	-0.323***	-0.163**	-0.095	-0.170***	-0.093	-0.166***	-0.175	-0.169***	-0.101	-0.089	0.041	-0.054
	(2.831)	(2.548)	(0.958)	(2.768)	(0.865)	(2.599)	(1.628)	(2.633)	(1.110)	(1.472)	(0.352)	(0.928)
Communal	-0.290***	-0.243***	-0.300***	-0.176***	-0.198**	-0.202***	-0.311***	-0.251***	-0.379***	-0.251***	-0.250**	-0.302***
	(3.119)	(4.580)	(3.612)	(3.457)	(2.223)	(3.892)	(3.514)	(4.810)	(5.072)	(5.016)	(2.489)	(6.275)
Transport	-0.080	-0.135***	-0.118	-0.084*	-0.158*	-0.115**	-0.139	-0.103*	-0.225***	-0.168***	-0.181*	-0.183***
	(0.891)	(2.736)	(1.503)	(1.780)	(1.780)	(2.150)	(1.402)	(1.920)	(3.106)	(3.415)	(1.714)	(3.859)
R-squared	0.124	0.234	0.313	0.388	0.194	0.343	0.496	0.587	0.159	0.287	0.409	0.497
N	1756	1756	1756	1756	1765	1765	1765	1765	1862	1862	1862	1862

Source: Author's calculations from EUSS.

Note: t-values in parentheses.

Table 4

Estimates of the Quantile Regressions, (Dependent Variable: Log-Wage) Continued

	2000				2003			
	$b(0.1)$	$b(0.5)$	$b(0.9)$	*OLS*	$b(0.1)$	$b(0.5)$	$b(0.9)$	*OLS*
Constant	4.920***	5.485***	5.988***	5.259***	4.897***	5.498***	5.947***	5.347***
	(22.080)	(40.324)	(25.366)	(37.477)	(22.349)	(49.225)	(23.414)	(40.908)
Age	0.047***	0.037***	0.033***	0.049***	0.046***	0.038***	0.038***	0.042***
	(4.340)	(5.575)	(2.795)	(7.112)	(4.279)	(7.062)	(3.134)	(6.721)
Age^2	-0.005***	-0.003**	-0.002	-0.004***	-0.005***	-0.003***	-0.003**	-0.004***
	(3.584)	(4.057)	(1.456)	(5.331)	(3.536)	(5.229)	(2.073)	(4.834)
Secondary	0.167***	0.227***	0.296***	0.241***	0.204***	0.222***	0.323***	0.266***
	(2.854)	(5.958)	(4.617)	(6.126)	(3.335)	(6.943)	(4.684)	(7.153)
University	0.603***	0.850***	0.940***	0.804***	0.491***	0.839***	0.999***	0.812***
	(5.900)	(14.381)	(10.674)	(13.154)	(4.976)	(17.318)	(9.501)	(14.397)
White-collar	0.349***	0.270***	0.507***	0.339***	0.177***	0.263***	0.292***	0.262***
	(5.902)	(7.062)	(8.189)	(8.588)	(3.109)	(8.730)	(4.492)	(7.455)
Professional	0.737***	0.815***	1.437***	0.972***	0.623***	0.733***	0.931***	0.740***
	(7.878)	(14.968)	(17.969)	(17.258)	(6.663)	(16.534)	(9.712)	(14.341)
Construction	-0.144*	0.075	0.009	-0.002	-0.005	0.016	-0.002	0.008
	(1.918)	(1.553)	(0.114)	(0.043)	(0.069)	(0.400)	(0.023)	(0.172)
Commerce	-0.303***	-0.086*	-0.040	-0.158***	-0.014	-0.100***	0.064	-0.026
	(4.378)	(1.888)	(0.536)	(3.364)	(0.193)	(2.716)	(0.793)	(0.608)
Financial	-0.099	0.041	0.005	-0.009	-0.031	0.078**	-0.014	0.072
	(1.330)	(0.851)	(0.061)	(0.184)	(0.410)	(2.026)	(0.176)	(1.609)
Personal	-0.200*	-0.115*	-0.109	-0.131*	0.042	-0.022	-0.164	-0.046
	(1.952)	(1.734)	(1.003)	(1.914)	(0.399)	(0.399)	(1.348)	(0.711)
Communal	-0.338***	-0.123**	-0.114	-0.182***	0.018	0.024	0.017	0.007
	(4.492)	(2.342)	(1.347)	(3.362)	(0.236)	(0.577)	(0.181)	(0.140)
Transport	-0.346***	-0.069	0.128	-0.122**	-0.163**	-0.060	-0.117	-0.098**
	(4.498)	(1.321)	(1.473)	(2.270)	(2.240)	(1.501)	(1.285)	(2.098)
R-squared	0.177	0.282	0.409	0.506	0.127	0.276	0.394	0.472
N	1673	1673	1673	1673	1765	1765	1765	1765

Source: Author's calculations from EUSS.

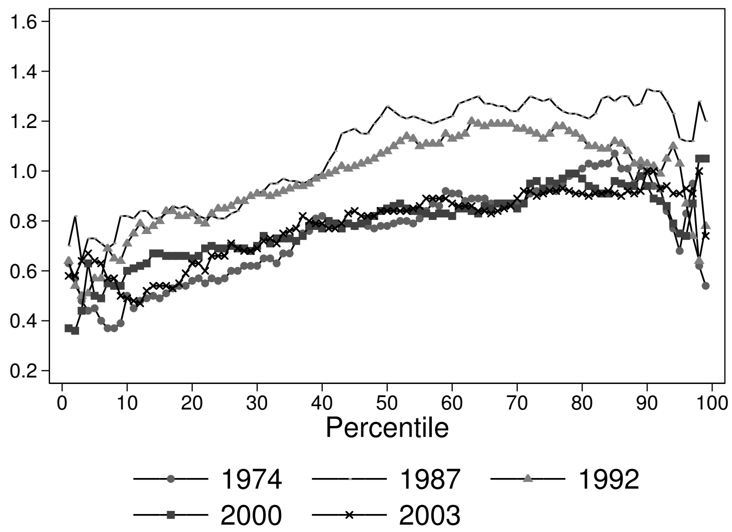
Note: t-values in parentheses.

It is also worth noting that the university education wage-premium clearly increased between 1974 and 1987, while it was virtually constant for secondary education. This implied that the payoff to an individual that had completed university education, compared to have completed secondary education, increased from 22% (0.386-0.160) in the first decile and 67% (1.021-0.347) in the ninth decile in 1974 to 65% and 100%, respectively, in 1987. Moreover, the university wage-premium declined in 1992, 2000, and 2003 in the lower part while it was fairly stable in the upper part of the wage distribution. For secondary education on the other hand, the wage-premium increased by 8.9% (0.258-0.169) in 1992 in the first decile and kept nearly at the same level in the nine decile, but in 2000 it returned to its level in the first two years.

Figure 4 reports in detail the parameter of University education at the different percentiles in five different years. For all five years the parameter increases as we move up in the male wage distribution. The parameter are higher, for almost all percentiles, in 1987 and 1992, than in 1974, 2000, and 2003. Comparing 1987 and 1992, while there are no clear differences until the 40:th percentile, for all the percentile above the 40:th the parameter is clearly higher in 1987. These results corroborate the findings of other studies (Beyer *et al.*, 1999) that point out a clear increase in the education wage-premium as Chilean trade became liberalized, although the percentage of university educated workers almost doubled between 1974 and 1987. The growing demand for skilled workers has already been mentioned as a reason for the widening of the skill premium, but the effect of the reduction of tariffs is also a factor that appears in the Chilean literature. The reduction of tariffs in the early 1970s induced a pressure on the prices of commodities intensive in unskilled workers, such as textiles, apparel, leather, and footwear, accounting for 18% of the employment in the manufacturing sector. In this way low-skill wages were under pressure, which helped to widen the skill premium (Beyer *et al.*, 1999).

Figure 4

University Parameter at the Different Percentiles for Years 1974, 1987, 1992, 2000, and 2003



Source: Author's calculation from Employment and Unemployment Survey for Santiago.

Another potential explanation is given by Pavcnik (2003) who suggests that the link between trade liberalization and the increased demand for skilled workers in Chile was investments. Trade liberalization reduced the relative price of imported machineries, materials, technology, and increased the competition from imported products. This resulted in increased investments by manufacturing plants, and thereby also increased the demand for skilled labor, as they are complementary.

The white-collar wage-premium increased in the ninth decile of the distribution until 1992 but declined after that, while in the first decile it fluctuated over the periods. For professionals, on the other hand, the wage-premium increased in 1987 and in 2000 in the upper part as well as in the lower part of the wage distribution.

The omitted sector in the regressions is manufacturing, so the parameters represent the marginal effect of moving a worker with given characteristics from manufacturing to another sector on the wage rate. One of the clearest patterns is that in the upper part of the wage distribution the number of significant industry parameters declined from three in 1974 to zero in 2003. On the other hand, in the lower part of the distribution it increased from three to five in 2000 but declined to one in 2003. For commerce, this parameter in the lower part of the distribution was significant in 1974, 1987, 1992, and 2000 but was significantly larger in the later three years when it ranged between -0.230 and -0.376 compared with -0.197 in 1974. This indicates that cross-industry wage differences have been more important among low paid workers than among highly paid ones. As these differences almost disappeared in 2003, one might say that this variable is not longer an important source of inequality in Santiago for the individuals analyzed here.

6 Decomposition Results

The results of the inequality decomposition are presented in Table 5. The first column displays the variance of log-wages, and columns (3), (5), and (7) represent the first decile, median and ninth deciles, respectively. Comparing 1974 with 1987, the most striking result is that the major contributor to the inequality change was the price effect, which explains roughly 50% of the change, followed by the composition effect with a somewhat smaller contribution. All the changes were found to be significant in the Bootstrap analysis.

Table 5

Results of the Decomposition of Inequality Measures, 1974-2003

	Variance (1)	(95% CI) (2)	Q(0.10) (3)	(95% CI) (4)	Q(0.50) (5)	(95% CI) (6)	Q(0.90) (7)	(95% CI) (8)
				1974-1987				
Total change	0.342*	(0.265;0.426)	0.265*	(0.185;0.328)	0.334*	(0.283;0.389)	0.839*	(0.700;1.030)
Price effect	0.188*	(0.054;0.280)	0.225*	(0.153;0.311)	0.270*	(0.204;0.353)	0.512*	(0.371;0.693)
Composition effect	0.129*	(0.030;0.208)	0.053	(-0.001;0.110)	0.090*	(0.034;0.161)	0.332*	(0.160;0.526)
Residual	0.026		-0.013		-0.026		-0.006	
Change only occupation	0.045	(-0.037;0.113)	0.016	(-0.049;0.064)	0.034	(-0.016;0.089)	0.097	(-0.074;0.252)
Change only education	0.171*	(0.104;0.242)	0.068*	(0.004;0.135)	0.152*	(0.087;0.210)	0.464*	(0.305;0.636)
Change only age	0.010	(-0.054;0.084)	0.002	(-0.047;0.056)	0.013	(-0.033;0.061)	0.030	(-0.118;0.204)
Change only industry	0.050	(-0.022;0.134)	-0.017	(-0.064;0.043)	0.020	(-0.038;0.079)	0.083	(-0.116;0.272)
				1987-1992				
Total Change	-0.272*	(-0.352;-0.192)	0.512*	(0.456;0.597)	0.435*	(0.371;0.526)	0.092	(-0.090;0.298)
Price effect	-0.215*	(-0.314;-0.075)	0.491*	(0.427;0.571)	0.398*	(0.318;0.503)	0.020	(-0.018;0.437)
Composition effect	-0.047	(-0.102;0.017)	0.030	(-0.009;0.074)	0.008	(-0.050;0.073)	-0.082	(-0.207;0.133)
Residual	-0.011		-0.010		0.029		-0.028	
Change only occupation	-0.007	(-0.061;0.042)	-0.002	(-0.040;0.055)	-0.004	(-0.045;0.042)	-0.021	(-0.165;0.137)
Change only education	-0.054	(-0.111;0.015)	0.008	(-0.030;0.058)	-0.006	(-0.047;0.050)	-0.117	(-0.238;0.067)
Change only age	-0.004	(-0.053;0.054)	0.006	(-0.038;0.049)	0.001	(-0.049;0.042)	-0.009	(-0.147;0.106)
Change only industry	-0.021	(-0.064;0.025)	-0.003	(-0.046;0.048)	-0.018	(-0.054;0.029)	-0.055	(-0.138;0.044)
				1992-2000				
Total Change	0.107*	(0.029;0.191)	0.422*	(0.373;0.471)	0.389*	(0.310;0.474)	0.435*	(0.270;0.547)
Price effect	0.066	(-0.042;0.178)	0.334*	(0.256;0.404)	0.352*	(0.292;0.424)	0.341*	(0.131;0.538)
Composition effect	0.037	(-0.042;0.102)	0.032	(-0.022;0.095)	0.059*	(0.019;0.111)	0.111	(-0.029;0.282)
Residual	0.003		0.057		-0.022		-0.017	
Change only occupation	0.029	(-0.028;0.104)	0.005	(-0.039;0.066)	0.014	(-0.029;0.065)	0.055	(-0.072;0.016)
Change only education	0.048	(-0.012;0.115)	0.023	(-0.041;0.074)	0.050*	(0.008;0.090)	0.110	(-0.004;0.222)
Change only age	0.003	(-0.057;0.058)	0.028	(-0.013;0.079)	0.026	(-0.016;0.070)	0.037	(-0.046;0.141)
Change only industry	0.031	(-0.026;0.103)	0.001	(-0.052;0.052)	0.020	(-0.017;0.060)	0.056	(-0.053;0.174)
				2000-2003				
Total Change	-0.128*	(-0.198;-0.060)	0.053	(-0.005;0.090)	0.023	(-0.049;0.093)	-0.116	(-0.231;0.011)
Price effect	-0.114*	(-0.216;-0.012)	0.070	(-0.006;0.170)	-0.002	(-0.072;0.084)	-0.093	(-0.277;0.072)
Composition effect	-0.003	(-0.060;0.058)	0.017	(-0.025;0.065)	0.017	(-0.022;0.066)	0.015	(-0.097;0.167)
Residual	-0.011		-0.035		0.008		-0.039	
Change only occupation	-0.011	(-0.063;0.056)	0.003	(-0.036;0.052)	0.003	(-0.038;0.044)	-0.006	(-0.108;0.112)
Change only education	0.007	(-0.038;0.049)	0.011	(-0.029;0.050)	0.020	(-0.020;0.065)	0.033	(-0.072;0.143)
Change only age	0.001	(-0.042;0.054)	0.015	(-0.026;0.051)	0.017	(-0.029;0.065)	0.020	(-0.083;0.145)
Change only industry	0.008	(-0.036;0.061)	0.006	(-0.037;0.046)	0.012	(-0.023;0.049)	0.036	(-0.069;0.129)

Sources: Author's calculations from EUSS.

Note: The numbers are calculated using only male salaried workers from EUSS surveys and 100 Bootstrap replications. $*$ Represent that the component is significant. *Variance, Q(0.10), Q(0.50)*, and *Q(0.90)* represent the variance of log-wages, the first decile, median, and ninth decile, respectively of the log-wage distribution. *(95% CI)* represents the 95% confidence interval calculated using a Bootstrap technique.

The results of the quintile regressions suggest that the price effect is most likely the result of the increased university education wage-premium, the professional occupation wage-premium, and probably the marked decline of the commerce sector parameter. The role of the rate of return to education has been already highlighted in other studies (Contreras, 2002); the new piece of information is that the change in the structure of workers contributed to the clear increase in inequality during this period. However, also in this contribution education appears as an important variable since this contribution was large in size and highly significant. Notice that when I calculate $I(f^*(w_{87})) - I(f^*(B_{87}; E_{74}))$ the education wage-premium and all other individual characteristics are the same in both distributions, so the change in inequality is only attributed to the increase in the level of education. The result is that the higher proportion of highly educated, and highly paid workers in 1987, compared with 1974, generates a large increase in the between education inequality. One also knows that the inequality in the right tail of the wage distribution increased. This implies that the counterfactual distribution $f^*(B_{87}; E_{74})$ contains a lower fraction of high-educated workers who belong to the part of the distribution where the dispersion increased between 1974 and 1987. In other words, the large effect of the change in the structure of education is explained by the fact that a larger proportion of workers became highly paid in 1987 and that the dispersion of income among highly paid increased between 1974 and 1987.

The price effect between these years is positive across the different quintiles but doubles in the ninth decile compared with the other two. Since zero is not included in the 95% confidence interval, these components are significant. Besides, the composition effect is much larger in the upper part of the distribution, where it accounts for 60% of the change compared with less than 30% in the other two.

Of these changes, the major contributor was the change in the educational structure which had a large, positive, and significant impact on wages across the quintiles. The effect is larger in the upper part of the distribution (0.46 in the ninth decile versus 0.06 in the first). The impact of occupation and industry are disproportionately high in this part of the wage distribution, but of minor intensity, and non-significant, if compared to education. These figures imply that the total effect was a clear increased polarization of the labor market with strong improvements in the top of the wage distribution and a poor improvement for low-wage workers.

The decline in inequality between 1987 and 1992 was mainly a result of the change in prices, which this time was large, and significant, in the lower part of the distribution and most likely reflects the sharp (28%) increase in the minimum wage between 1989 and 1993, the increase of the secondary education wage-premium in the first decile, and the disproportional decline of the university parameter in the ninth decile. It is also noteworthy that the small change of $Q(0.90)$ was the result of two opposite effects, namely the positive price effect and the negative composition effect. However, non of the components in the decomposition of this decile were significant. All together contributing to the recovery of the wages in the bottom and mid of the distribution and a reduction of inequality.

In the following period, the inequality change was more attenuated in comparison with the previous periods because the similar size of the growth across the different deciles. This time the change in the variance of log-wages and in all the deciles were significant. The inequality increasing price effect was the most important once again, explaining almost the totality of the change, followed by the composition effect. However, the composition effect was larger in the right tail of the wage distribution, where it accounts for 25% of the change compared with 7%-15% in the first and fifth deciles, respectively. Nevertheless, only the change in the median was significant according to the Bootstrap analysis. The effect of occupation, education, and industry was more important for the ninth decile. In the last period, inequality declined despite the small change in the first and fifth deciles since the ninth declined (mainly because of the inequality decreasing price effect) generating a 17% drop of the variance of log-wages. In contrast to the period 1992-2000, in the 2000-2003 period only the total change in the variance of log-wages was significant.

7 Conclusions

This paper has applied a quintile regression approach to analyze the wage inequality changes between the years 1974, 1987, 1992, 2000, and 2003 in Santiago's labor market. Inequality increased vary significantly between 1974 and 1987, but declined fast over a few years. The period 1992-2003 is characterized by an inverted-U pattern of earnings inequality. Especially important were the changes in inequality in the upper part of the distribution, where it followed a similar trend to that of the overall inequality, being more or less constant in the lower part of the wage distribution.

The inequality decomposition applied in this paper has been used in several studies recently (Nguyen *et al.*, 2007; Albrecht *et al.*, 2003), but thus far not on research with Chilean data. Calculating counterfactual distributions of wages I am able to analyze the effect of the change of the parameters of the wage equations (the price effect) and the effect of the change in the distribution of individual characteristics such as age, educational level, occupation, and sector of employment (the composition effect). In essence, this method is inspired in Oaxaca-Blinder decompositions but here is not applied to the mean of the distribution but to inequality measures.

The price effect emerges as the most important source of inequality change, but has its largest and significant effect when inequality declined between 1987 and 1992. Underlying the remarkable inequality increase between 1974 and 1987 is a combination of a price effect (which accounts for about 50% of the change) and a composition effect (accounting for 37% of the change). The composition effect had a particularly relevant role in the large increase of the ninth decile, generating an increased polarization of the labor market with a strong improvement at the top of the wage distribution but a poor improvement for low wage workers. Moreover, the change in the level of education was the only variable that had a significant effect in all deciles.

The pattern in the following period was completely different when the first and the fifth deciles increased the most, as a result of the effect of prices, generating a recovery of the wages at the bottom and in the middle of the distribution which in turn reduced inequality. During the 1990s, the growth of male earnings was very similar across the different deciles and even the price effect was of the same size. The composition effect, on the other hand, was larger in the upper part of the distribution (0,11) in relation to the lower part (0,059 and 0,032). The result was an inequality increase between 1992 and 2000, though of a considerably smaller size than between 1974 and 1987.

In summary, the results lead one to the conclusion that fluctuations of the price of individual characteristics are the main source of fluctuations in wage inequality during the last decades. The change in the structure of individual characteristics was important only comparing 1974 with 1987.

An interesting further application of the approach used in this paper is include female workers in the analysis and evaluate the impact that the increased female participation, of which a large proportion is highly educated, have had on the inequality of the last decades.

References

Author, D. H., L. F. Katz, and M. S. Kearney (2005): "Rising Wage Inequality: The Role of Composition and Prices," *Harvard Institute of Economic Research*, Discussion Paper Number 2096.

Albrecht, J., A. Björklund, and S. Vroman (2003): "Is There a Glass Ceiling in Sweden?," *Journal of Labor Economics*, 21(1), pp. 145-177.

Athanasopoulos, G., and F. Vahid (2003): "Statistical Inference and Changes in Income Inequality in Australia," *The Economic Record*, 79, pp. 412-424.

Beyer, H., P. Rojas, and R. Vergara (1999): "Trade Liberalization and Wage Inequality," *Journal of Development Economics*, 59, pp. 103-123.

Chile Social and Economic Indicators 1960-2000 (2002): *Banco Central de Chile*, Santiago, Chile.

Contreras, D. (2002): "Explaining Wage Inequality in Chile: Does Education Really Matter?" *The Journal of Income Distribution*, 11(1), pp. 53-76.

Cortázar, R. (1997): "Chile: The Evolution and Reform of the Labor Market" Chapert 8 in *Labor Markets in Latin America: Combining Social Protection with Market Flexibility*, S. Edwards and N. C. Lustig (eds.), Brookings Institution Press.

Edwards, S., and A. C. Edwards (2000): "Economic Reforms and Labour Markets: Policy Issues and Lessons from Chile," *Economic Policy*, 15:30, pp. 181-230.

Ffrench-Davis, R. (2005): "Economic Reforms in Chile, From Dictatorship to Democracy," Michigan: The University of Michigan Press.

Gottschalk, P., and T. M. Smeeding (1997): "Cross-National Comparisons of Earnings and Income Inequality," *Journal of Economic Literature*, Vol. XXXV, pp. 633-687.

Krongkaew, M. (1994): "Income Distribution in East Asian Developing Countries: An Update," *Asian-Pacific Economic Literature*, 8(2), pp. 58-73.

Larrañaga, O. (2001): "Distribución de Ingresos en Chile: 1958-2001, Department of Economics," Universidad de Chile, Working Paper 178.

Machado, J. A. F., and J. Mata (2004): "Counterfactual Decomposition of Changes in Wage Distributions Using Quintile Regression," *Journal of Applied Econometrics*, 19, pp. 1-21.

Melly, B. (2005): "Decomposition of Differences in Distribution Using Quantile Regression," *Labour Economics*, 12, pp. 577-590.

Meller, P., R. O'Ryan, and A. Solimano (1996): "Growth, Equity, and the Environment en Chile: Issues and Evidence," *World Development*, 24:2, pp. 255-272.

Mills, J. A., and S. Zandvakili (1997): "Statistical Inference Via Bootstrapping for Measures of Inequality," *Journal of Applied Econometrics*, 12, pp. 133-150.

Montenegro, C. E. (1998): "The Structure of Wages in Chile 1960-1996: An Application of Quintile Regression," *Estudios de Economia*, 25(1), pp. 71-98.

Moran, T. M. (2006): "Statistical Inference for Measures of Inequality With a Cross-National Bootstrap Application," *Sociological Methods and Research*, 34(3), pp. 296-333.

Nguyen, B. T., J. W. Albrecht, S. B. Vroman, and M. D. Westbrook (2007): "A Quintile Regression Decomposition of Urban-Rural Inequality in Vietnam," *Journal of Development Economics*, 83, pp. 466-490.

Palma, A. (2005): "Income Inequality in Chile, Trends and Explanations: A Survey," *Working Paper*, Department of Economics, Göteborg University.

Pavcnik, N. (2003): "What Explains Skill Upgrading in Less Developed Countries?" *Journal of Development Economics*, 71(2), pp. 311-328.

Psacharopoulos, G., S. Morley, A. Fiszbein, H. Lee, and W. Wood (1995): "Poverty and Income Inequality in Latin America During the 1980s," *Review of Income and Wealth*, 41(3), pp. 245-264.

Robbins, D. (1994): "Relative Wage Structure in Chile, 1957-1992: Changes in the Structure of Demand for Schooling," *Estudios de Economía*, 21, pp. 51-78.

Zin, R. H. M. (2005): "Income Distribution in East Asian Developing Countries: Recent Trends," *Asian-Pacific Economic Literature*, 19(2), pp. 36-54.

Appendix

Table A1

Definition of Variables.

	Definition
Wage	Monthly wages and salaries of male workers/ hours worked.
Age	Age.
Age^2	Age squared and divided by 10.
Primary	Dummy = 1 if no education or primary education, otherwise 0.
Secondary	Dummy = 1 if post-primary non-university education, otherwise 0.
University	Dummy = 1 if university education, otherwise 0.
Blue-collar	Dummy = 1 if blue-collar occupation, otherwise 0.
White-collar	Dummy = 1 if white-collar occupation, otherwise 0.
Professional	Dummy = 1 if professional occupation, otherwise 0.
Manufacturing	Dummy = 1 if Manufacturing, otherwise 0.
Construction	Dummy = 1 if Construction, otherwise 0.
Commerce	Dummy = 1 if Commerce, otherwise 0.
Financial	Dummy = 1 if Financial services, otherwise 0.
Personal	Dummy = 1 if Personal services, otherwise 0.
Communal	Dummy = 1 if Communal services, otherwise 0.
Transport	Dummy = 1 if Transport, otherwise 0.

Table A2

Changes in the Variance of Log-Wages (ΔVar), 1974-2003

		1987	1992	2000	2003
1974	ΔVar	0.358*	0.074*	0.177*	0.053
	(95% CI)	(0.282;0.446)	(0.022;0.158)	(0.100;0.259)	(-0.015;0.129)
1987	ΔVar		-0.284*	-0.181*	-0.305*
	(95% CI)		(-0.352;-0.300)	(-0.268;-0.093)	(-0.386;-0.235)
1992	ΔVar			0.103*	-0.021
	(95% CI)			(0.048;0.181)	(-0.086;0.039)
2000	ΔVar				-0.123*
	(95% CI)				(-0.202;-0.056)

Source: Author's own calculations from EUSS.
Notes: Calculated using 100 Bootstrap replications.

Table A3

Changes in the Gini Coefficient of Wages ($\Delta Gini$), 1974-2003

		1987	1992	2000	2003
1974	$\Delta Gini$	0.129*	0.036*	0.087*	0.019
	(95% CI)	(0.092;0.167)	(0.008;0.065)	(0.059;0.124)	(-0.013;0.056)
1987	$\Delta Gini$		-0.093*	-0.042*	-0.110*
	(95% CI)		(-0.119;-0.060)	(-0.071;-0.005)	(-0.140;-0.085)
1992	$\Delta Gini$			0.051*	-0.017
	(95% CI)			(0.024;0.082)	(-0.040;0.009)
2000	$\Delta Gini$				-0.069*
	(95% CI)				(-0.099;-0.036)

Source: Author's own calculations from EUSS.
Notes: Calculated using 100 Bootstrap replications.

Table A4

Changes in the Theil Inequality Index of Wages ($\Delta Theil$), 1974-2003

		1987	1992	2000	2003
1974	$\Delta Theil$	0.267*	0.046	0.209*	0.023
	(95% CI)	(0.198;0.358)	(-0.011;0.123)	(0.109;0.209)	(-0.055;0.092)
1987	$\Delta Theil$		-0.221*	-0.058	-0.244*
	(95% CI)		(-0.298;-0.153)	(-0.176;0.053)	(-0.333;-0.180)
1992	$\Delta Theil$			0.163*	-0.022
	(95% CI)			(0.068;0.247)	(-0.097;0.054)
2000	$\Delta Theil$				-0.186*
	(95% CI)				(-0.272;-0.082)

Source: Author's own calculations from EUSS.
Notes: Calculated using 100 Bootstrap replications.