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CALM AFTER THE STORMS:

INCOME DISTRIBUTION IN CHILE, 1987-1994

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Abstract: After rising in the 1960s, falling in the early 1970s, and rising again from the mid-seventies to the mid-eighties, Chilean inequality seems to have stabilized since around 1987. After the 'stormy' period of economic and political reforms of the 1970s and 1980s, no statistically significant Lorenz dominance results could be detected since 1987. Scalar measures of inequality confirm this picture of stability, but suggest a slight change in the shape of the density function, with some compression at the bottom being 'compensated for' by a stretching at the top. As inequality remained broadly stable, sustained economic growth led to substantial poverty reduction, according to a range of measures, and with respect to three different poverty lines. Poverty mixed stochastic dominance tests confirm this result. All of these findings are robust to different choices of equivalence scales. An examination of the factors underlying these trends suggests that an equilibrium has (for the moment) been reached between rising demand for and supply of skills, where the former is associated with technological progress, and the latter with expansions in education. Chile's trading patterns might well be tangential to its recent distributional dynamics.

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1. Introduction

Two 'storms' have recently raged over the distribution of income in Chile. The first, and most important, was caused by a series of structural reforms of the economy - which started in 1974 and were largely completed by the late eighties - and by successive changes in political regime - which had important implications, among other things, for the regulation of labor markets. The economic reforms included trade liberalization, privatization of state-owned assets, deregulation of various markets, and reforms in the structure of taxes, subsidies and benefits. They have been extensively discussed elsewhere, and are well beyond the scope of this paper. See Edwards and Edwards (1987) and Scott (1996) for excellent summaries. The political changes were fundamentally the military coup d'etat of 1973, which installed General Pinochet as President, and the restoration of democracy in 1990, with the election of President Aylwin.

The second storm, closer to the 'tea cup' variety, has raged in academic and policy circles, as the effects of the "Chilean model" on poverty and inequality were hotly debated. This storm also had two separable components which, for convenience, I will name "Chile versus Stolper-Samuelson" and "See, All is Not Well After All". The former component originates from the finding that, as Chile liberalized its trade regime, the ratio of skilled workers' wages to those of the unskilled rose, rather than fell (Robbins, 1994). If the Chilean and the world economies could be well approximated by a model where unskilled and skilled labor were the only two factors of production, and the other Hecksher-Ohlin assumptions held (notably constant technology and no non-tradable goods), then this finding would violate the predictions of the Stolper-Samuelson theorem, which establishes the links between goods and factor prices in a Hecksher-Ohlin world. If Chile - like other developing countries - had relatively abundant supplies of unskilled labor, opening up to trade should increase the returns to this factor relative to those to skilled labor - the opposite of what Prof. Robbins found.

The second strand ("See, All is Not Well After All") draws on Robbins's findings, but percolates beyond academia to Chilean politics and society more broadly. The Concertacion governments of Presidents Aylwin and Frei have made "Growth with Equity" their paramount objective. A failure to promote equity - indeed the observation of actual increases in inequality - can be seized upon as evidence of their failure to deliver on their stated objectives. Income distribution statistics became increasingly important in the Chilean political and social debate of the mid-1990s, to the point where the Catholic Church's *Conferencia Episcopal de Chile* issued an open letter in January 1996, entitled "Is Chile an Equitable Country?". Its concluding section opened by stating that "...the current distribution of income in our country should be cause for scandal among Christians...The distance between the rich and the poor has grown in Chile in an alarming fashion. A solidary country can not accept this reality. Neither can a modern country tolerate these differences." (Comision Nacional de Justicia y Paz, 1996, p.24).

Indeed, although there was no single *national* household survey which collected detailed information on incomes from all sources with a regular periodicity prior to 1985, the consensus is that inequality in Chile did rise substantially throughout the 1960s, 1970s and 1980s (except for a brief decline in 1970-73). This information is largely based on the long time-series of the Universidad de Chile Household Survey data, which covers only Greater Santiago. Riveros (1983) was the first to note the rising trend in the Gini coefficient, focusing on the period from 1958 to 1982. This was corroborated by Robbins (1994), and by Montenegro (1996), for wage incomes, although both authors found that the trend changed in 1990. More recently, using the only *national* data available, Londono and Szekely (1997) confirmed these findings for total household incomes, across the entire country. Data points satisfying these requirements prior to 1990 are only available for 1971, 1980 and 1989. Londono and Szekely (1997) report the Gini coefficient rising over those three years from 0.47 to 0.53 and 0.59. They too found a reversal beginning in 1990, with the Gini falling to 0.57 in 1994.

Nevertheless - although Robbins (1994) finds a decline in skilled/unskilled wage ratio from 1990 to 1992; Montenegro (1996) finds that the wage Gini for greater Santiago falls from 0.57 in 1987 to 0.46 in 1996; and Londono and Szekely report the aforementioned small decline in the national total income Gini between 1989 and 1994 - there does not yet seem to be a consensus on the post-1990 part of the story. Robbins (1995) himself plays down the post-1990 decline, attributing it to the rise in numbers of university graduates from 'low-quality' private universities. The Bishops who wrote the aforementioned 1996 letter clearly also doubt that there has been an improvement since the turn of the decade. Having become accustomed to three decades of rising inequality, it is as if Chilean society (and some foreign economists) refuse to believe that a reversal was possible, and wonder what economic mechanisms might underlie it. Meller (1996), suggested that a reduction in wage disparities might be due, after all, to Stolper-Samuelson at work. Robbins (1996) and Wood (1997) remained unconvinced.

In part, room for this controversy is generated by the absence of a thorough, definitive analysis of the best available data since 1987. Such an approach should be based on as solid a treatment of the data as possible, taking issues such as regional price variations and differences in family composition into account. It should also overcome the inherent ambiguity of inequality analysis based on scalar measures: income share ratios might fall, while Ginis rise; coefficients of variation might suggest increasing dispersion, while the mean log deviation suggests declines. And it should deal explicitly with the statistical nature of inequality analysis: some of the changes observed, particularly in the 1990s, have been so small that one wonders whether they are statistically significant.

This paper aims to fill that gap, and to contribute to the debate on the dynamics of personal income distribution in Chile by establishing which facts and conclusions are indeed empirically robust. It provides a comprehensive description of the levels of and changes in poverty and inequality in Chile from 1987 to 1994, drawing primarily on a detailed analysis of four household survey microdata sets - the CASEN surveys of 1987, 1990, 1992 and 1994 - which are discussed briefly in Section 2 below, and described in more detail in

Appendix 1. We address the ambiguity of scalar welfare, poverty and inequality measurement by relying on stochastic dominance techniques. The paper incorporates the need for statistical testing of hypotheses about changes in income distribution, by applying the Howes (1993) intersection-union test for statistical significance of stochastic dominance. It also addresses the issue of robustness with respect to the assumptions made about different needs across households and economies of scale within them, by presenting all results for equivalised income, as well as for per capita income. Finally, we present a set of static and dynamic inequality decompositions, to shed some light on the possible economic determinants of our findings, and comment on how these might relate to previous work on Chilean income distribution in an open-trade regime.

The structure of the paper is as follows. The next section briefly discusses the concepts analyzed in this study, describes the data set on which the analysis is based, and some of our methodological assumptions. Section 3 presents the basic results on inequality and welfare, by means of scalar measures, decile means and shares, and stochastic dominance. Section 4 discusses the evolution of the poverty indicators, after presenting the rationale for our choice of poverty lines. Section 5 delves behind the statistics, and seeks to shed some light on the factors that can help explain the levels and changes in inequality, through both static and dynamic decomposition analyses. It also suggests how the results may be reconciled with earlier findings in the "Chile versus Stolper-Samuelson" debate. Section 6 concludes.

2. Concepts, Data and Methodology

In studying the evolution of the Chilean distribution of income during 1987-1994, this paper presents results concerning three distinct, but related concepts: social welfare, poverty and inequality. Social welfare is perhaps the vaguest of the three concepts: it seeks to capture the level of well-being of a society or population. Due to information constraints, it is usually proxied in household survey studies by income or consumption expenditure[2]. Many

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[2] Among those two, there are two reasons why expenditure is usually the preferred welfare indicator. First, if capital markets work at all, it is a better proxy for permanent income. Second, there is a growing

have argued for more encompassing measures, which would include a value for leisure or other non-monetary dimensions of quality of life, such as the availability of public goods, whether environmental in nature or not.[3] While the inclusion of these dimensions would be ideal in principle, we are constrained in practice - by data availability and reliability - to relying on either consumption expenditure or income.

Since the CASEN surveys, widely regarded in Chile as the best available sources of information on households since their creation in 1985, use income as the welfare concept surveyed, so will this paper. Hence, the social welfare measures we will be concerned with are non-decreasing functions of current income, and are symmetric (not discriminating between different recipients along any dimension other than income). We will not present values for any specific social welfare measures in this study, relying instead on the theorems of stochastic dominance discussed below to investigate changes in social welfare.

In seeking to measure welfare, whose current income should one consider? The income distributions conventionally studied in the Chilean literature have been household income per capita per household distributions; which is to say that the income unit was household income per capita, and the recipient unit was the household. This causes a number of inconveniences, such as having smaller shares of the population in higher deciles[4] than in lower ones (as household size declines with household income per capita in Chile, see MIDEPLAN, 1992, p.34), or having a headcount poverty measure which is less than the

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consensus among practitioners that household survey income data is less reliable than expenditure data, due to a number of problems related to mis-reporting and mis-measurement. See Deaton (1995) and Chauduri and Ravallion (1994) for discussions. Chile's last *available* expenditure survey, however, is the *Encuesta de Presupuestos Familiares* of 1987/88. Apart from being old and incapable of providing any information on changes during the recent high-growth period, it only covered the Metropolitan Region of Santiago.

[3] Becker (1965) proposed a 'full income' concept which included an imputed value for leisure. Sen (1981) proposed a concept of entitlements which aimed to capture some of the benefits of public goods, in addition to the value of income.

[4] Strictly speaking, a decile is a separator. There are nine deciles, and decile i is the income which separates tenth i from tenth i+1 in the distribution. Like other quantiles (e.g. percentiles or quintiles), the word for decile is widely misused to refer to the actual tenth (or hundredth, or fifth) of the distribution. Rather than confuse the reader by departing from the usual misnomer, decile will be used to mean tenth in this Report, as will its finer and coarser analogues.

proportion of the population (in persons) which is poor. We therefore adopt the individual as the recipient unit throughout: all our distributions are vectors of individuals. The income they are imputed is one of two income concepts we adopt: household income per capita, to allow for some comparisons with previous studies and for ease of direct interpretation; and household income per equivalent adult, to take account of differences in needs arising from distinct household sizes and compositions. A discussion of the specific equivalence scale used is included below.

Poverty can be thought of as the negative of some welfare function, over a censored distribution. It is negative, in the sense that a rise in poverty, everything else constant, is a decline in welfare. More importantly: poverty measures are defined over a censored distribution in the sense that they measure the welfare of those below a certain threshold in the overall distribution.[5] That threshold is the poverty line which, in this case as in many others, is defined as an absolute income level (fixed over time in real terms) which is judged just sufficient to provide an individual with minimum nutritional and other requirements. Details are discussed in Section 4.

Inequality, like welfare, is defined over complete distributions of an indicator. Unlike welfare, however, it is independent from the mean of those distributions, concerning itself only with their second moment. Inequality measures are defined over mean-normalized income distributions, and are generally required to satisfy the Pigou-Dalton transfer principle, which demands the measure to rise (or at least not fall) in response to a mean-preserving spread.

There are, of course, many different such inequality measures, which may quite validly rank the same two distributions in opposite ways, if they are more sensitive to distances in different parts of the distribution. The best remedy to this ambiguity is to rely on stochastic dominance, which is discussed below, but we will also present four different inequality

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[5] Most poverty measures, including all those used in this study, satisfy the focus axiom, which requires them to be invariant with respect to any change in income levels above the poverty line.

measures, to provide information on changes from a number of different ‘perspectives’. We will present values for the Gini coefficient (which is most sensitive to incomes in the middle of the distribution), the mean log deviation (which is most sensitive to incomes in the bottom of the distribution), the Theil index (whose sensitivity is constant across the distribution) and a transform of the coefficient of variation (which is most sensitive to incomes at the top of the distribution). Their formulae are given in Section 3.

Let us now turn briefly to the data sets of the Caracterización Socioeconómica Nacional (CASEN), for the years of 1987, 1990, 1992 and 1994, on which the analysis below is based.[6] The CASEN is a nationally and regionally representative[7] household survey carried out by the Chilean Ministry of Planning (MIDEPLAN), through the Department of Economics of the Universidad de Chile, with the dual objectives of generating a reliable portrait of socioeconomic conditions across the country, and of monitoring the incidence and effectiveness of the government’s social programs and expenditures. With these ends, questions are asked pertaining both to the household and to the individuals within the household. Topics covered include demographics; characteristics of the dwelling; access to utilities and public services; educational attainment (if currently enrolled, detailed questions are asked about the school, method of education financing, benefits, etc.); health conditions; health insurance; health services used and benefits received; occupation and employment; and incomes. Income questions are designed to permit the distinction between labor incomes in cash, labor incomes in kind (agricultural and non-agricultural), income from capital, rental income, imputed rent, employment-related transfers (such as occupational, invalidity or widow’s pensions.) and entitlement transfers (such as the basic pensions PASIS or the family allowance SUF). Details on the sampling methodology employed by Universidad de Chile, as well as on the adjustments made to the raw data by CEPAL, are provided in Appendix 1.

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[6] A first CASEN survey was conducted in 1985, but it is widely reported to be less reliable, of inferior quality than and less comparable with the subsequent ones.

[7] Representativeness at the level of the comuna has also become increasingly widespread, as sample sizes have increased over time, but can not be assumed for all comunas.

The income variable from the CASEN used as the welfare indicator in this study is total household income, further adjusted in two ways. First, the income vector was deflated by a regional price index, with Santiago as the base location. Traditionally, nominal incomes have only been deflated by a common national consumer price index, taking no account of regional variations in price levels which, as Table 1 indicates, can in some cases be considerable. This is reinforced by Chile's geography, with extreme southern and northern regions having substantially higher average prices than those closer to Santiago. Our regional price deflation is based on the only source of prices outside Santiago, the National Statistical Institute's (INE) *Anuario de Precios* survey of 16 cities.[8] Given the variation in price levels from year to year, we used an average of the index from 1985 to 1994; the values are given in Table 1 below. Regionally adjusted incomes were then deflated over time by the national consumer price index for November (the survey month) of the relevant years, as given in CEPAL (1995, p. 24). All income values reported in this chapter, therefore, are expressed in 1994 Santiago pesos.

Table 1: Regional Prices

Region	**Price Index**
I	1.2163
II	1.1668
III	1.1112
IV	1.0859
V	1.0420
VI	1.0269
VII	1.0136
VIII	1.0441
IX	1.0500
X	1.0383
XI	1.0856
XII	1.2208
XIII (RM)	1.0000

The second adjustment was to introduce the concept of household income per equivalent adult, through the adoption of an equivalence scale. This was done to capture the changes in measured inequality and poverty which arise out of taking into account the differences in needs between households with different compositions (e.g. four adults as opposed to a couple with two small children), as well as economies of scale which arise from sharing fixed housing or other costs. There are a number of different approaches to deriving an equivalence scale,

[8] There are no systematic surveys of prices in rural areas anywhere in Chile. Though in the past rural prices have been assumed to be lower than in urban areas, we have found no justification for this arbitrary mark-down, and generalized urban prices to the whole region.

and there is no single accepted dominant method. Rather than attaching excessive importance to the specific values of our chosen coefficients, we sought to provide a reasonably reliable alternative to the per capita income concept, which is well known to constitute an extreme assumption in terms of differences in needs (none), as well as in terms of (the absence of) economies of scale.

Our chosen scale is a revised version of the equivalence scale for Chile, calculated by Contreras (1995), using the Rothbarth adult goods method. Contreras's scale was estimated excluding all households with a single adult from the sample, and taking two-adult households as the reference type. He found that adult good expenditures were restored to the childless couple level when incomes for families with one child in the age categories below was raised by the percentage amount indicated[9]:

Table 2: Proportional Compensating Variations for Children in Chile

Child age	No child	0-4 yrs	5-10 yrs	11-15 yrs
Cost increase	0%	15%	20%	40%

Since we must also cover those households made up by a single individual, and in order to take into account some economies of scale within the household, we assume the cost of a single person to be 60% of that of a couple (roughly equivalent to saying that the second adult costs 70% of the cost of the first adult[10]). Our equivalence scale is therefore:

$Y_i = X_i / M_i$, with $M_i = 1.2 + 0.8\,(N_{aa} + N_{11\text{-}15}) + 0.4\,N_{5\text{-}10} + 0.3\,N_{0\text{-}4}$

where: N_{aa} is the number of additional adults in the household;

$N_{11\text{-}15}$ is the number of children aged 11-15 (inc) in the household;

$N_{5\text{-}10}$ is the number of children aged 5-10 (inc) in the household;

$N_{0\text{-}4}$ is the number of children aged 0-4 (inc) in the household.

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9 The entries in Table 2 are a bottom-line approximation. While they capture the results at an appropriate level of confidence, they do not do justice to the complexity of the estimation method, and do not take into account the different standard errors associated with different age categories. See Contreras (1995) for these and other details.

10 This is a common assumption, adopted for instance in the construction of the OECD equivalence scale (see OECD, 1982).

Note that this has maintained households with two adults as the reference group. Their household income will be divided by two. An additional child in the 11-15 age category 'costs' an extra 40%, as before. An additional child in the 5-10 category costs an extra 20%, as before. An additional child in the 0-4 age category costs an extra 15%, as before. The second adult accounts for 40% of the couple's total costs.

By introducing the equivalence scale and the regional price adjustments directly on to the incomes to be analyzed, consistency is ensured in the assumptions underlying the inequality and the poverty analyses. Also, since all incomes are effectively expressed in 1994 Santiago pesos, there will be no need for regional poverty lines. Lines expressed in 1994 Santiago pesos will be the appropriate comparators for all incomes. Similarly, there will be no need for poverty lines for different household types; it will suffice to compare the household incomes per equivalent adult with an individual adult poverty line.[11] The advantage of this approach over introducing those concepts through different poverty lines, in addition to simplicity, is that the inequality analysis would not then - as it does now - incorporate regional price and equivalence scale adjustments.

In the next section, the analysis of inequality and welfare changes relies on equivalised household income, based on the equivalence scale described above. In order both to preserve comparability with previous studies and to demonstrate the robustness of the main results to the adoption of the scale, the analysis is replicated for household incomes per capita, in Appendix 2.

3. Levels and Changes in Chilean Inequality and Social Welfare

Before presenting the detailed results, it is worthwhile highlighting the broad picture which arises from the data, and which can be characterized by three 'stylized facts'. First, the

[11] Though the absolute values of the poverty measures do, of course, depend on the household reference type chosen in defining the scale.

entire distribution function has been shifting to the right over time, with people in the same relative ranks earning higher incomes in later years. Though this is not strictly the case for every segment of the distribution in every year, it is the general case.[12] This is clearly the result of economic growth. Second, the dispersion of the distribution seems to have remained broadly stable as it moved to the right over this period. If anything, there is some indication of a slight reduction in overall inequality, though there are no significant and unambiguous changes in inequality between any of the years surveyed by CASEN.[13] This would suggest that the benefits of economic growth have been distributed in a pattern roughly similar to that of the existing income distribution. Third, to the extent that there were any discernible changes in the shape of the density function, within the broad context of stability, these appear to have been a slight compression in the lower tail, and a slight increase in dispersion in the upper tail: inequality among the poor fell, while inequality among the very rich, and between them and those just poorer than them, seems to have increased.

As stated in the previous section, the four scalar inequality measures used in this study are the Gini Coefficient, the mean log deviation, the Theil index and a transform of the coefficient of variation. In their formulae below, we use the following standard notation: y_i is the income of individual i, $i \in (1, 2, ..., n)$, n is the number of individuals in a given distribution, and $\mu(\mathbf{y})$ is the arithmetic mean of the distribution.

The Gini is given by $G = \frac{1}{2n^2 m(y)} \sum_{i=1}^{n} \sum_{j=1}^{n} \left| y_i - y_j \right|$.

The other three measures are all members of the Generalized Entropy class of inequality indices, which satisfy a number of desirable properties, such as symmetry, population

[12] The first order welfare dominance results reported below specify the instances when it was actually true for the entire distribution.

[13] In the sense that there are no statistically significant Lorenz dominances between any two years in our sample. See below.

replication, scale invariance and decomposability. See Cowell (1995) for details. The general formula for the parametric class is given by:

$$E(a) = \frac{1}{a^2 - a}\left[\frac{1}{n}\sum_{i=1}^{n}\left(\frac{y_i}{m(y)}\right)^a - 1\right] \tag{1}$$

Using l'Hopital's rule, one can obtain E(0), the mean log deviation:

$$E(0) = \frac{1}{n}\sum_{i=1}^{n}\log\left(\frac{m(y)}{y_i}\right)$$

Similarly, the Theil index corresponds to E(1), which is given by:

$$E(1) = \frac{1}{n}\sum_{i=1}^{n}\frac{y_i}{m(y)}\ \log\left(\frac{y_i}{m(y)}\right)$$

The fourth measure we have employed is E(2), which can be expressed as:

$$E(2) = \frac{1}{2nm(y)^2}\sum_{i=1}^{n}\left(y_i - m(y)\right)^2$$

Table 3 below lists mean and median incomes, as well as these four inequality measures, for the household income per equivalent adult distribution in each of the four years we analyze.

Table 3: Descriptive Statistics: Monthly Household Incomes per Equivalent Adult

	1987	**1990**	**1992**	**1994**
Mean Income	67,232	75,007	90,797	93,981
Median	36,265	42,455	50,212	53,196
Gini	0.5468	0.5322	0.5362	0.5298
E(0)	0.5266	0.4945	0.4891	0.4846
E(1)	0.6053	0.5842	0.6151	0.5858
E(2)	1.3007	1.3992	1.5050	1.5634

The impact of economic growth can be seen immediately, through the sharp upward trend in mean and median incomes. The large differences between the mean and the median, which persist over the period, are an indication of the skewness of the distribution. The four measures of inequality confirm the high level of inequality in Chile, by international standards.[14]

In terms of temporal evolution, the Gini coefficient, which is especially sensitive neither to the top nor to the bottom of the distribution, changes very little over the period. Though there is a slight decline in its value, it is unlikely to be statistically significant. The Theil index (E(1)) behaves slightly differently, displaying a sine pattern. It does not suggest a strong trend in either direction either. The other two measures are suggestive of the small changes referred to above: the mean log deviation (E(0)), which is particularly sensitive to low incomes, falls a little more markedly than the Gini throughout. E(2), a transform of the coefficient of variation which picks up differences in the upper tail with greater weight, rises monotonically from the beginning to the end of the period. Overall, there appears to be limited change in inequality, but with greater distances at the top compensating for smaller distances at the bottom. Tables 4 and 5 below allow us to investigate a more disaggregated picture, by listing decile means and shares for per capita household incomes.

Table 4: Decile Mean Incomes: Household Incomes per Equivalent Adult (1994 Santiago Pesos)

	1987	**1990**	**1992**	**1994**
Decile 1	9036.39	10405.88	13811.73	13468.71
Decile 2	16205.31	19245.45	23615.58	24126.88
Decile 3	21285.65	24970.63	30687.34	31567.66
Decile 4	26700.95	31421.56	37740.56	39265.97
Decile 5	32816.62	38564.88	45765.64	48300.51
Decile 6	40599.89	47079.28	55930.66	59518.87

[14] While the average Gini coefficient for Chile in this period - with respect to distributions of household incomes per capita (see Appendix 2) for comparability - was 0.5539, a recently compiled international inequality database indicates that the average Gini in the 1980s (1990s) was 0.3323 (0.3375) in industrialized countries and high-income developing countries; 0.2501 (0.2894) in Eastern Europe; 0.3501 (0.3188) in South Asia; 0.3870 (0.3809) in East Asia and the Pacific; 0.4045 (0.3803) in the Middle East and North Africa; 0.4346 (0.4695) in Sub-Saharan Africa and 0.4975 (0.4931) in Latin America and the Caribbean. See Deininger and Squire (1996). Not only is Chile in quite a different inequality league from industrialized countries, or indeed from those in Asia, but it also has a Gini coefficient above the Latin American average.

Decile 7	51487.53	59378.09	70167.22	74523.33
Decile 8	68820.84	77958.43	92232.57	99180.40
Decile 9	105610.50	116361.20	134597.00	148112.80
Decile 10	299725.50	324647.70	403411.90	402088.70
Top Percentile	808409.30	926408.60	1241728.00	1180673.00

Table 5: Decile Income Shares: Household Incomes per Equivalent Adult (%)				
	1987	**1990**	**1992**	**1994**
Decile 1	1.34	1.39	1.52	1.43
Decile 2	2.41	2.57	2.60	2.57
Decile 3	3.17	3.33	3.38	3.36
Decile 4	3.97	4.19	4.16	4.18
Decile 5	4.88	5.14	5.04	5.14
Decile 6	6.04	6.28	6.16	6.33
Decile 7	7.66	7.92	7.73	7.93
Decile 8	10.24	10.39	10.16	10.55
Decile 9	15.71	15.51	14.82	15.76
Decile 10	44.58	43.28	44.43	42.73
Top Percentile	12.02	12.35	13.68	12.41

Table 4 offers remarkable confirmation of the gains from economic growth to all deciles of the Chilean income distribution. In fact, every single decile has seen its average income rise in every sub-period, with only two exceptions: the bottom and the top deciles in 1994. This sustained increase in real incomes across the distribution, over a period of seven years, is an achievement most countries would be proud of.

As for the exceptions, the fall in mean income for the first decile in 1994 has been the subject of considerable debate in Chile since the data first became available. This study confirms that the decline also took place in terms of household incomes per equivalent adult. There seems to be little question that it was due, at least in part, to the decline in the overall rate of GDP growth, which was 11.8% in the second semester of 1992, and 4.3% in the same period of 1994. In particular, this cyclical deceleration - brought about largely by contractionary monetary policy aimed at curbing inflationary pressures - caused an increase in unemployment from 4.8% to 6.5%[15]. The unemployment rate was much higher in the poorest quintile, rising there from 18% to 22% (See Cowan and De Gregorio, 1996). Since labor earnings are such an important component of the incomes of the poor, the reduction in the demand for unskilled labor which is behind the above employment figures is bound to have contributed to the recorded decline in their overall incomes.[16] It means, of course, that the sustained increases in social welfare achieved from 1987 to 1992 will not continue unambiguously into 1994, despite continued GDP growth. Similarly, for some (low) poverty lines there would be an increase in poverty from 1992 to 1994, as indeed was the case for some poverty measures with respect to the indigence line reported in the next section.

Table 5 abstracts from means and allows us to focus on inequality, as depicted by decile shares. The overall impression is once again of a stable (mean-normalized) distribution, with changes in shares being generally small in proportion to the shares themselves. Nevertheless, there is some evidence of the trend for a compression at the bottom and increased dispersion at the top, certainly until 1992. For the first three years in the sample, decile shares for the bottom three deciles rise, while those for deciles 8 and 9 fall. The top decile is trendless, but there is some indication that incomes at the very top are climbing faster than others, with the share of the richest 1% of the population rising over the period.

1994, however, does represent a break in this trend. The decline in mean incomes at the top and bottom must imply a reduction in shares for deciles 1 and 10. Since the overall mean continued to grow, we in fact see a reduction in the shares of the first three deciles. The gainers were the middle-classes, broadly defined, of deciles 4 to 9. It must be emphasized, however, that despite an absolute loss in incomes in the bottom decile - which does have implications for poverty and welfare - it would be wrong to conclude, as many commentators have, that inequality increased unambiguously from 1992 to 1994. This is

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[15] In the three months ending November (the CASEN survey month) of both years, according to Cowan and De Gregorio (1996).

because decile shares provide merely a (somewhat) disaggregated view of the distribution, rather than an accurate yardstick of inequality. There is no Lorenz dominance of 1992 over 1994 and, indeed, three of our four measures actually fall in that interval. From the evidence presented in this study, one could only claim that inequality (defined so as to be consistent with the Pigou-Dalton transfer principle) has worsened if he chose E(2) as the only measure that mattered.

In closing this section, we turn to the results of our stochastic dominance analysis. As the preceding discussion illustrates, inequality comparisons between different distributions depend on the specific measures being employed, and ambiguities are often inevitable. The concept of stochastic dominance, which originates from the analysis of financial risk, was introduced to the field of income distribution analysis to help establish when unambiguous comparisons of inequality or social welfare are possible.

Distribution A displays first-order stochastic dominance over distribution B if its cumulative distribution function $F_A(y)$ lies nowhere above and at least somewhere below that of B, $F_B(y)$. For any income level y, fewer people earn less than it in A than in B. If that is the case, a theorem due to Saposnik (1981) establishes that any social welfare function which is increasing in income will record higher levels of welfare in A than in B.

Distribution A displays second order stochastic dominance over B if its deficit function (the integral of the distribution function $G(y_k) = \int_0^{y_k} F(y)dy$) lies nowhere above (and somewhere below) that of B. It is a weaker concept than its first order analogue, and is in fact implied by it. Shorrocks (1983) has shown that if it holds, any social welfare function that is increasing *and* concave in income will record higher levels of social welfare in A than in B.

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[16] See also Beyer (1995), for an interesting discussion of the patterns of employment and labor force participation in the first quintile.

The dominance criteria described above are alternative concepts suitable for comparing welfare. Inequality requires us to abstract from the mean, and concentrate on the dispersion of the distribution. For this purpose, mean-normalized second-order dominance - also known as Lorenz dominance - is the appropriate concept. Distribution A is said to Lorenz dominate distribution B if the Lorenz curve associated with it lies nowhere below, and at least somewhere above that associated with B. A Lorenz curve is a mean-normalized integral of the inverse of a distribution function: $L(p) = \frac{1}{m(y)}\int_0^p F^{-1}(p)dp$. In other words, it plots the share of income accruing to the bottom p% of the population, against p. For a Lorenz curve (A) to lie everywhere above another (B) means that in A, the poorest p% of the population receive a greater share of the income than in B, for every p. Atkinson (1970) has shown that if it holds, inequality in A is lower than in B according to any inequality measure that satisfies the Pigou-Dalton transfer axiom.

Table 6 below presents the results of these three types of dominance comparison among the four years for which we have CASEN data, for both income concepts. The comparisons were carried out first at the percentile level of aggregation, and then checked for the completely disaggregated sample, with its statistical significance tested according to the endogenous bounds method of Howes (1993). A letter F, S, or L in cell (i, j) in Table 6 indicates that year i respectively first-order, second-order or Lorenz dominates year j. The letter was inserted when dominance was found at the percentile level. An asterisk indicates that the dominance was statistically significant at the 95% confidence level over a range greater than or equal to 99% in Howes's endogenous bounds test for the complete sample, and hence that the hypothesis of no population dominance can be rejected at that level. This analysis was carried out for both the household income per capita and the household income per equivalent adult distributions, and the results (including their significance tests) were identical in every case except one. The exception was the (statistically insignificant) Lorenz dominance of 1994 over 1990, which was found only for the per capita income distribution, and is therefore entered in brackets in the table below. All Lorenz comparisons for equivalised income display crossings.

Table 6: Welfare and Inequality Stochastic Dominance Comparisons				
	1987	**1990**	**1992**	**1994**
1987				
1990				
1992	**F*, S***	**S***		
1994	**F*, S***	**F*, S*, (L)**		

Two observations can immediately be made: first, there are a number of significant welfare dominance results below the diagonal, indicating that welfare rose unambiguously from some earlier to some later years. Second, with the exception of the statistically insignificant case of 1994 over 1990 for per capita income, there are no cases of Lorenz dominance. This suggests that, with the possible exception of an improvement between 1990 and 1994, inequality comparisons between the years in this period are ambiguous, and will depend on the specific measure used. This confirms the results presented in Table 3 above.

Much more can be said in terms of the evolution of social welfare. Since our measure of welfare depends entirely on income, as discussed in the introduction, one would expect rapid economic growth to have a powerful impact. Nevertheless, the dominance results are interesting because they tell us something about the distribution of the gains from growth across households. Rapid growth between 1987 and 1990, for instance, was not sufficient to lead to unambiguous welfare gains, because the poorest one or two percent of the population were worse off in 1990. Gains above the second or third percentile meant that there was still a rise in the mean income of the first decile, as reported earlier in this section. The very disaggregated nature of dominance analysis allows us to capture finer changes. This loss to the very poorest people in Chile at the end of the last decade meant that welfare functions very sensitive to their circumstances would not have shown an increase in social welfare since 1987, despite the substantial increase in incomes elsewhere in the distribution.

Growth from 1990 to 1992 did not seem to have this perverse effect at the bottom of the distribution. Income rises across every percentile ensured that 1992 and 1994 both first-order dominate 1987. Both years also second-order dominate 1990.[17] 1994 and 1992, however, can not be ranked, whether by the first or second order criterion. This is due to a decline in incomes for those below the eighth or ninth percentile (which, on this occasion, was sufficient to lower the mean income of the first decile, as seen in Table 4 above), which co-existed with gains for all other social groups. This welfare loss to the very poor means that social welfare can not unambiguously be said to have risen over the last two years in our sample. The loss was not sufficient, however, to outweigh gains to those at the bottom of the distribution since 1990: 1994 does display first-order dominance over both 1987 and 1990.

These results add rigor to our earlier analysis, while broadly supporting its findings. The first fundamental feature of the period is economic growth, which led to welfare dominance of the last two years over the first two. On two occasions, however, economic growth failed to improve measured living standards for the most vulnerable people in society: from 1987 to 1990 and, more famously, from 1992 to 1994. For those pairs of years, no such unambiguous welfare comparison is possible.

The second fundamental feature - the relative stability in the dispersion of the distribution, but with a slight compression at the bottom and a stretching at the top - is also compatible with the absence of significant Lorenz dominance results. Such changes in the shape of the density function, reducing distances in one part of the distribution while simultaneously increasing them elsewhere, is exactly what causes different inequality indices to rank distributions in opposite ways. While this section has emphasized that the evidence on inequality is not sufficiently clear to identify any real tendency in either direction, so that the most appropriate description is one of broad stability or inconclusive changes, the (insignificant) Lorenz dominance of 1994 over 1990 for the per capita income distribution

[17] 1992 does not first-order dominate 1990 because of a crossing above the 99th percentile. Except for the very rich, everyone (in an 'anonymous' sense) was better off in 1992 than in 1990.

does confirm that, if one were really pushed to suggest a direction for change in inequality over the period, it would more likely be a reduction.

4. The Evolution of Poverty

The high GDP growth rates achieved by Chile over the last ten years have undeniably contributed to a considerable reduction in poverty from the relatively high levels of the mid-1980s. This section presents detailed results on the changes between 1987 and 1994, relying on our adjusted data set and comparing the numbers for household per capita income with those for household income per equivalent adult. Though the general downward trend confirms previous findings (see, e.g. Larranaga, 1994, and Contreras, 1995), these new numbers reflect our adjustments, such as the incorporation of regional price differences, the new equivalence scale adopted and the improved treatment of domestic servants. Before presenting the specific results, we briefly discuss the derivation of the poverty lines, with respect to which all of the measures must be understood.

We use three poverty lines in this study, all of them expressed in 1994 Santiago pesos per month: an indigence line set at P$15,050, a lower-bound poverty line set at P$30,100, and an upper-bound poverty line equal to P$34,164. The first two are the official indigence and poverty lines widely used in Chile. The incomes with which they are compared, however, differ from most earlier studies in that they have been ‘converted’ to 1994 Santiago pesos, using as ‘exchange rate’ the regional price index in Table 1 and the November CPI. The derivation of the upper-bound line is explained below.

All three lines are absolute poverty lines, and derive from a standard food basket specified by CEPAL. The basket is chosen so as to provide 2,187 Kcal per person per day, the national average caloric requirement, which is obtained from the demographic characteristics of the population and from the FAO/WHO recommended caloric intakes for different age and gender groups. The specific commodity composition of the basket is based on the actual consumption patterns of a reference group chosen by CEPAL, which is the third quintile by expenditures in the Household Expenditure Survey of 1987/88. It is valued at average prices for November 1994 in Santiago. The monthly cost of this standard CEPAL food basket has traditionally been used in Chile as an 'indigence' (or extreme poverty) line, separating the hard core of the poor - those whose current monthly incomes are insufficient even for the purchase of a minimum diet - from the rest of society. In 1994, this amount was P$15,050, which is therefore reported below as the indigence line.

In accordance with international practice, however, that is deemed too strict a criterion to identify the poor. There are, after all, other basic expenditures in addition to food which everyone must make, such as shelter, clothing and public transport. A standard methodology is then applied to arrive at a sensible poverty line: the cost of the food basket is multiplied by the inverse of the food share in total expenditure (the Engel coefficient) for some suitable reference group. Based on the estimates of the Engel coefficient for the lower quintiles in the Chilean expenditure distribution, which are reported in Table 7 below, we decided to adopt the standard value of 0.5, which implies a doubling of the indigence line to arrive at the poverty line.[18] However, though Table 7 is reassuring in that the coefficients vary little from total household expenditure to expenditure per capita (suggesting robustness with respect to the equivalence scale adopted), they are substantially lower for the concept of per capita income including imputed rent. Since this is the welfare concept which we have adopted for a large part of our analysis[19], it was felt that its implications for poverty measurement could not be ignored. Weighing the relevant coefficients for the first and second quintiles by 0.8 and 0.2 respectively, one arrives at an Engel coefficient of approximately 0.44. Applying its inverse (2.27) to the cost of the food basket yields our upper-bound poverty line of P$34,164.

Table 7: Engel Coefficients: Households ordered by:

Quintile	**Total Expenditure**	**Exp. per capita**	**Income per capita (inc. imputed rent)**

[18] This has the important advantage of allowing some comparability with the findings of earlier works, most of which have used this line.

[19] And since the household income per equivalent adult concept also includes imputed rents.

1	0.530	0.539	0.451
2	0.494	0.489	0.397

Source: CEPAL (1996), p. 31, Table 13.

We also use the same three lines discussed above when computing poverty measures using the vector of real household incomes per equivalent adult. Given our choice of two-adult households as the reference household type, this means that the per capita poverty line is unchanged for that household type. As with any equivalence scale designed to take account of different relative costs of children and of economies of scale, ours implies different per capita poverty lines for household types other than the reference. This reflects the re-ranking of households which the scale inherently causes. Table 8 below sets out what our equivalence scale implies in terms of household and per capita poverty lines.

Table 8: Implied Household and Per Capita Poverty Lines

Household Type (1)	**Equivalence Factor (2)**	**Household Poverty Line [(2) x 30,100] = (3)**	**Implied Per Capita Poverty Line [(3) / #persons] = (4)**
Single Adult	1.2	36,120	36,120
Couple	2.0	60,200	30,100
Couple + Child (0-4)	2.3	69,230	23,077
Couple + Child (5-10)	2.4	72,240	24,080
Couple + Child(11-15)	2.8	84,280	28,093
Couple + Child (0 - 4) + Child (11 - 15)	3.1	93,310	23,328

Let us now turn to the poverty measures. For a given poverty line (z), different poverty indices can be defined, each aggregating information on the living standards of those below the poverty line in different ways. We work with three of the commonest measures, all of which can be expressed as members of the following parametric class, proposed by Foster, Greer and Thorbecke (1984):

$$P_a = \frac{1}{n}\sum_{i=1}^{n}\left[\max\left(\frac{z-y_i}{z},0\right)\right]^a$$

where y_i is the income assigned to the ith individual (of which there are n). These measures are rather intuitive. As is well known, when $\alpha = 0$, P simplifies to p/n, the headcount index. When $\alpha = 1$, we have the normalized poverty deficit and $\alpha = 2$ yields the FGT(2) measure, which incorporates some convexity to the distances between incomes and the poverty line, and is hence sensitive to inequality among the poor. Table 9 below lists the values of each of these measures for the whole country, in each relevant year, for the distribution of household per capita incomes. Each index is listed for each of the three poverty lines discussed above.

Table 9: Poverty Measures: Household Incomes per Capita

	1987	**1990**	**1992**	**1994**
Indigence Line:	P$ 15,050			
Headcount	0.2209	0.1646	0.1403	0.0996
Poverty Deficit	0.0756	0.0561	0.0325	0.0336
FGT (2)	0.0382	0.0295	0.0172	0.0184
Poverty Line L:	P$ 30,100			
Headcount	0.5137	0.4427	0.3603	0.3386
Poverty Deficit	0.2274	0.1838	0.1329	0.1269
FGT (2)	0.1299	0.1017	0.0681	0.0663
Poverty Line H:	P$ 34,164			
Headcount	0.5679	0.5002	0.4206	0.3940
Poverty Deficit	0.2647	0.2181	0.1637	0.1554
FGT (2)	0.1560	0.1240	0.0861	0.0831

Table 10 is analogous to Table 9, and lists the values of the same measures for the whole country, in each relevant year, for the distribution of household incomes per equivalent adult.

Table 10: Poverty Measures: Household Incomes per Equivalent Adult

	1987	**1990**	**1992**	**1994**

Indigence Line:	P$ 15,050			
Headcount	0.1268	0.0894	0.0474	0.0511
Poverty Deficit	0.0412	0.0311	0.0174	0.0192
FGT (2)	0.0213	0.0176	0.0108	0.0118
Poverty Line L:	P$ 30,100			
Headcount	0.4069	0.3306	0.2418	0.2308
Poverty Deficit	0.1568	0.1196	0.0776	0.0762
FGT (2)	0.0822	0.0614	0.0376	0.0382
Poverty Line H:	P$ 34,164			
Headcount	0.4726	0.3889	0.3000	0.2852
Poverty Deficit	0.1905	0.1483	0.1006	0.0978
FGT (2)	0.1028	0.0777	0.0491	0.0492

According to all three measures, there has undoubtedly been a remarkable reduction in both poverty and extreme poverty from 1987 to 1994. Poverty in Chile was quite high in the mid-1980s, in the aftermath of the serious recession of 1982-84. 51-57% of the population lived in poverty in 1987, according to the per capita income concept, or 41-47% if one relies on the income per equivalent adult concept, which takes better account of household needs and economies of scale.[20] By 1994, the figures were 34-39% for per capita incomes, or 23-29% for equivalised incomes. [21] The incidence of indigence fell from 22% to 10% by the per capita concept, or 13% to 5% by the equivalised concept.

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[20] These poverty ranges refer to the values with respect to the lower-bound and upper-bound poverty lines.

[21] The poverty headcount figures for the per capita distribution are higher than those reported by various earlier studies (see Haindl, 1996, for a survey). This is due primarily to the following methodological corrections which have been made: (a) Some earlier studies report as a headcount the proportion of households below the poverty line. We report the proportion of individuals below the poverty line. Since households tend to be larger among the poor, this increases the figure. (b) Incorporating regional price differences raises poverty since Santiago prices are the lowest in the country: deflating incomes elsewhere to take this into account reduces their real incomes. (c) Because we did not apply the arbitrary 0.66 factor to reduce prices in rural areas (see Appendix 1), we did not lower poverty lines for rural areas. We believe that there was no justification for that practice, particularly in light of the fact that reported figures for own consumption from household production among the rural poor did not appear to be unrealistic. The empirically baseless reduction of rural prices by an arbitrary factor might well have been contributing to an under-estimation of rural poverty. (d) Finally, our inclusion of domestic servants as individuals with their own incomes - rather than those of their employers - might have added some individuals to the ranks of the poor. See Appendix 1.

As others have pointed out before, the reductions were largest in the years of faster growth, from 1987 to 1992, and smallest in the relatively more sluggish years of 1992-94. In fact, though all headcount indices for the per capita income distribution record poverty reductions in that latter period, as do the other two measures for poverty, it is noteworthy that the poverty deficit and the FGT(2) measures for indigence actually record slight increases in extreme poverty between 1992 and 1994. The picture is even more severe for the equivalised income distribution where, in addition to increases in P_1 and P_2, the indigence headcount also rises, suggesting that the number of those living in extreme poverty increased from 1992 to 1994, despite continued economic growth. Furthermore, FGT(2) also rises (marginally) for the other two poverty lines as well.

This is a result of the decline in real incomes within the first decile of the income distribution, which was discussed in the previous section. Clearly, the closer a poverty line is to that decile, the likelier it is to record an increase, whereas more generous lines still record declines as a result of income gains to people in the second and third deciles. But even for those lines, poverty measures which are more sensitive to large distances between incomes and the poverty line (i.e. that place greater weight on greater destitution) - such as FGT(2) - are liable to pick up the losses at the very bottom and have them outweigh gains closer to the upper poverty lines.

This picture of considerable reductions in poverty throughout the period, albeit with some ambiguity between 1992 and 1994, is confirmed by stochastic dominance analysis. It has been shown that if a distribution A displays poverty mixed dominance (PMD(z^-, z^+)) over a distribution B, then any poverty measure which is decreasing in income, satisfies the focus axiom and the transfer axiom (in situations where a crossing of the poverty line does not occur), will indicate that poverty is lower in A than in B, for any poverty line between z^- and z^+.[22] This class of poverty measures includes all members of the Foster-Greer-Thorbecke P_α class, and is therefore particularly appropriate for this study. Poverty mixed dominance essentially requires that distribution A display second-order dominance over B from zero to the lower bound poverty line (z^-), and first-order dominance from z^- to z^+. Table 11 below, which is analogous to Table 6, presents the results for Chile, with z^- set at the indigence line and z^+ set at the upper-bound poverty line. The letter P in cell (i, j) indicates that year i dominates year j. As before, dominance was checked for both per capita income and income per equivalent adult; on this occasion, both concepts yield exactly the same results, so that there are no entries in brackets.

Table 11: Poverty (Mixed) Stochastic Dominance Comparisons

	1987	**1990**	**1992**	**1994**
1987				
1990				
1992	P	P		
1994	P	P		

These results reveal that there was unambiguously less poverty in 1992 and in 1994 than in either 1987 or 1990, whether poverty is measured by the headcount, the poverty deficit or indeed any of a host of other sensible poverty measures, and for any poverty line set anywhere between P$15,050 and P$34,164 per month in 1994 Santiago pesos. This sort of unambiguous poverty reduction, quite independent of the specific measure used and valid for such a large range of poverty lines, is not common. Its achievement confirms the widely held view that Chile has made substantial strides in the fight against poverty during the last decade.[23]

Yet, there is also confirmation of two sub-periods in which growth did not lead to unambiguous poverty reductions: from 1987 to 1990 and from 1992 to 1994. On both occasions, although the headcount for the headline poverty line indicates a reduction in the

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[22] See Howes (1993) for a discussion, and Ferreira and Litchfield (1996) for an application to Brazil.

[23] All but one of the positive dominance results reported in the above table follow directly from the first order-dominances reported in Table 9. The table contains new information only for the cases where there was no dominance, as well as for the dominance of 1992 over 1990, the distribution functions of which clearly do not cross between z^- and z^+.

number of poor people, there were income losses in the lowest percentiles of the distribution. These losses imply that some poverty measures in the wide class defined above would have indicated increases in poverty for at least some of the poverty lines in the covered range. Indeed, in the second sub-period this was the case for all three indigence measures reported in Table 10, for the equivalised income distribution. Still, the dominance of 1994 over both 1987 and 1990 indicates that the losses to some of the poor in the last two years were at least not sufficient to outweigh the gains made between 1990 and 1992.

Overall, there is no question that Chile's growth and social policies[24] were tremendously successful in reducing the incidence, intensity and inequality of poverty between 1987 and 1992, with the poverty deficits being roughly halved across all three poverty lines (a little less for the per capita incomes, a little more for the equivalised incomes). The recent reversal in the performance of the poorest of the poor, between 1992 and 1994, does however provide a cautionary signal against complacency.

5. *Static and Dynamic Decompositions of Inequality*

This paper has focused on a detailed description of the levels of and changes in poverty, inequality and welfare in Chile during the last decade. This section briefly discusses some of the structural factors which may explain Chilean inequality. The analysis relies on two sets of inequality decompositions: the first group are static decompositions, which aim to separate inequality measures into within-group and between-group components, and the second set are dynamic decompositions, which shed some light on the nature of changes in inequality over time. The income concept used for all decompositions is household income per equivalent adult.

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[24] Larra⊙aga (1994) has decomposed the changes in poverty in Chile between 1987 and 1992 into a growth and a redistribution component, using a methodology due to Datt and Ravallion (1992). While he found that some 80% of the reduction could be explained by the effects of growth, some of the changes were also due to a redistribution effect, which may very well have followed - at least in part - from the government's social policies and expenditures.

The basic idea behind static inequality decompositions is that household and personal characteristics, such as education, gender, occupation and regional location, are important determinants of household income. If that is the case, then at least part of the value of any given inequality measure must reflect inequality between people with different educational levels, occupations, genders, and so on. This inequality is referred to as the "between-group" component. But for any such partition of the population - whether by region, occupational sector or any other attribute - some inequality will also exist among people in the same sub-groups; this is the "within-group" component. While inferring causality from such decompositions requires considerable circumspection, especially when the attribute defining the partition is variable (such as education), one often refers to the between-group component as the share of inequality 'explained' by that particular attribute, while the within-group component is 'unexplained' or residual.

Although many common inequality measures are not decomposable across such partitions in a meaningful way, all members of the Generalized Entropy class can be decomposed in a very simple form. Let $\Pi(k)$ be a partition of the population into k subgroups, indexed by j. Let $\mu(y)_j$ be the mean income in subgroup j; $E(\alpha)_j$ be the inequality measured for the population in subgroup j; $f_j = n_j/n$ be the population share of subgroup j ; and $v_j = \frac{n_j m(y)_j}{nm(y)}$ be the income share of subgroup j. If we define the between-group component I_B as: $I_B = \frac{1}{a^2 - a}\left[\sum_{j=1}^{k} f_j \left(\frac{m(y)_j}{m(y)}\right)^a - 1\right]$ and the within group component I_W as: $I_W = \sum_{j=1}^{k} w_j E(a)_j$ where the weights are given by: $w_j = v_j^a f_j^{1-a}$, then Cowell and Jenkins (1995) show that overall inequality I can be written simply as $I = I_B + I_W$.[25]

[25] Cowell and Jenkins (1995) draw on earlier work by Bourguignon (1979), Cowell (1980) and Shorrocks (1980 and 1984).

The share of inequality explained by a given partition Π, for a specific Generalized Entropy measure I, which is reported in Table 12, is simply $R_B(\Pi) = I_B(\Pi)/I$. This 'explained' share depends not only on the attribute defining the partition, but also on the specific member of the Generalized Entropy class being decomposed. Below, we use the following household or personal attributes to define partitions: the region where the household lives; its urban/rural status; gender of head; age of head; education of head and occupation of head. The value of R_B for each of these partitions is given in Table 12 below for both the mean log deviation (E(0)) and the Theil index (E(1)). It constitutes a measure of the relative importance of the relevant attribute in explaining inequality, as measured by the specific index.

The regional partition divides the population into 13 subgroups, according to which of the country's 13 regions the household is located in. Naturally, the urban/rural partition divides the population into two subgroups, as does the partition based on gender of the household head. The partition by age of the head of the household creates six sub-groups: less than 25 years, 25 to 34 years, 35 to 44 years, 45 to 54 years, 55 to 64 years and 65 years or older. The education partition divides households into five subgroups, according to whether the head has less than four years of schooling, between four and seven years of schooling, between eight and eleven years of schooling, between twelve and fifteen years of schooling, or more than fifteen years of schooling. The occupational partition divides households into eleven groups, according to whether the head reports an occupation which fits into the following sectoral categories: outside labor force; unemployed; agriculture; mining; manufacturing; utilities (electricity, gas, water); construction; commerce; transport, storage or communications; financial and other 'entrepreneurial' services; social and personal services.

Table 12: Static Decomposition Results: R_B

Partition	**1987**		**1990**		**1992**		**1994**	
	E(0)	E(1)	E(0)	E(1)	E(0)	E(1)	E(0)	E(1)
Region	0.05	0.05	0.05	0.04	0.08	0.06	0.08	0.06

Urb/Rur	0.07	0.05	0.02	0.02	0.04	0.03	0.05	0.04
Gender	0.01	0.01	0.01	0.01	0.01	0.00	0.00	0.00
Age	0.01	0.01	0.01	0.01	0.02	0.01	0.02	0.02
Education	0.30	0.32	0.26	0.24	0.32	0.31	0.26	0.24
Occupation	0.10	0.10	0.09	0.09	0.07	0.06	0.10	0.08
Fine	0.58	0.58	0.55	0.55	0.56	0.55	0.55	0.54

Table 12 sheds considerable light on the relative importance of these personal and household attributes in explaining inequality. In interpreting it, one should bear in mind that each individual decomposition is unrelated to the others, so that the R_B values are *not* analogous to partial derivatives or regression coefficients in a multiple regression. Each partition does *not* control for the other attributes. The first finding is that regional location can not account for a very substantial share of observed inequality in Chile. Though its importance rose over the period, it started at 5% and never reached 10% of the total. While this suggests that broad regional income disparities, such as are observed in other countries (such as Brazil or Indonesia), are not a big issue in Chile, it should not be taken to mean that where a household lives is of no use as a targeting variable. Indeed it has been suggested that income disparities at the level of the comuna - rather than the region - can be very substantial (see World Bank, 1997). Geography may be related to inequality but, if it is, it is at a finer level of disaggregation than at the macro-regional level.

Inequality between urban and rural areas also fails to capture a very significant proportion of overall inequality, accounting for only 2 - 7% of the total. According to these decompositions, however, the least important personal attributes in explaining inequality are the gender and age of the household head. Let us take these in turn. The unimportance of gender (0-1% of the total over the entire period) is in line with international evidence for developing countries. Ferreira and Litchfield (1997) find a gender R_B of 0.00 for Brazil, in all three years investigated (1981, 1985 and 1990). Quisumbing et al (1995) used stochastic dominance techniques to investigate whether male-headed households fared better than female-headed ones in ten developing countries, and statistically rejected that hypothesis in most cases. Nevertheless, as these authors point out, these results refer only to inequality between *households* headed by males or females. It says nothing about labor-market disparities or intra-household distribution between males and females, and should not be construed as evidence relating to those important issues.[26]

The unimportance of age is perhaps more surprising. Life-cycle theories of human capital and income, and considerable evidence from labor market studies, suggest that earnings are significantly correlated with age, and one might expect this to show up in disparities between households headed by different age groups. The finding, which is similar to that of Ferreira and Litchfield (1997) for Brazil, suggests that households somehow dilute those earning differentials by age. Since the variable being decomposed is household income per equivalent adult, it is reasonable to suppose that the birth of children compensates for (at least part of) the increase in income which may accrue to a household head as s(he) ages. Through childbirth or otherwise, taking households as the unit and dividing its income by the number of equivalent adults offsets whatever income disparities exist between different age groups in the labor market: the decomposition does indicate that inequality between different age groups of households is not important.

The two most important attributes for explaining inequality in Chile are education and occupational sector of the household head. And of the two, education is clearly the crucial factor. Depending on the measure and on the year, between one-quarter and one-third of total inequality is accounted for by differences in the educational attainment of the household head, with mean incomes for each category rising markedly with schooling. Compared to this, even the 6 - 10% accounted for by occupational sector seems small, particularly in light of the fact that educational differences are not controlled for in the occupational decomposition. This result suggests that the current policy focus on education is not misplaced, and that continued improvements in educational attainment by the poor are

[26] See Agrawal and Walton (1996) for a survey of labor market gender issues in a number of developing countries.

likely to lead to reductions in overall inequality, in addition to their direct effect on poverty reduction.

Finally, the bottom row of Table 12 refers to the proportion of overall inequality which is explained by differences between the 5720 subgroups in a fine partition by all six attributes in the table, *combined together*. This proportion is remarkably consistent over time and across the two measures, remaining always in the 54-58% range. By international standards, this is a relatively high proportion of overall inequality accounted for by personal and household attributes (see Cowell and Jenkins, 1995, and Ferreira and Litchfield, 1997). Nevertheless, looked at in another way, these figures indicate that even when the population has been very finely partitioned, into nearly 6,000 subgroups - each of which consists of households that live in the same region and zone, and whose heads belong to the same gender, age, education and occupational categories - still some 42-46% of total income inequality is within-group, and can not be accounted for by differences between the subgroups. This is in line with the international experience that inequality is not only a complex variable to measure without ambiguity, but also a phenomenon whose magnitude is difficult to explain by the standard set of personal or household attributes.

Turning now to the *changes* in inequality, we rely on a decomposition of the mean log deviation (E(0)), due originally to Mookherjee and Shorrocks (1982), which separates a change in overall inequality into four components, again in relation to a partition by household attributes. Mookherjee and Shorrocks (1982) have shown that the overall change in E(0) can be well approximated by the following expression:

$$\Delta E(0) \cong \sum_{j=1}^{k} \overline{E(0)_j} \Delta f_j + \sum_{j=1}^{k} \left[\overline{l}_j - \overline{\log(l_j)} \right] \Delta f_j + \qquad \text{(a + b)}$$

$$\sum_{j=1}^{k} \left(\overline{v_j} - \overline{f_j} \right) \Delta \log\left(m(y)_j \right) + \qquad \text{(c)}$$

$$\sum_{j=1}^{k} \overline{f_j} \Delta E(0)_j \qquad \text{(d)}$$

where an overbar indicates a simple average of the values of the variable at t and at t+1; and $\lambda_j = \mu(y)_j/\mu(y)$ is the ratio of subgroup mean to overall mean.

The decomposition has four additive terms (a - d). The sum of a and b gives the population shares effect; c is the relative mean incomes effect, and d is the unexplained effect. Their values (after both sides are divided by E(0)) are entered in Table 13 below for the indicated partitions and time intervals.

Table 13: Dynamic Decomposition Results: 1987 - 1994

%ΔE(0)	-8.7			
Partition	**a**	**b**	**c**	**d**
Region	0.0	0.0	1.5	-10.2
Urb/Rur	0.7	-0.8	-1.3	-7.3
Gender	0.0	0.0	-0.8	-7.9
Age	-0.4	-0.1	1.4	-9.5
Education	2.1	4.2	-10.8	-4.1
Occupation	3.3	1.3	-2.1	-11.1

a and **b** measure the ‘population shares’ effect;
c measures the ‘mean incomes’ effect;
d measures the ‘unexplained’ effect.

In interpreting the table, we note that the overall change in E(0) was negative in the period considered (see Section 3),[27] and that a negative sign in any component indicates that that

[27] The overall change in E(0) differs slightly from the change that can be derived from Table 3, due to small sample differences resulting from the elimination of households with missing values for

term contributed towards the decline in inequality as measured by the mean log deviation. The first remarkable observation is that in all cases but education, the decline in overall inequality is not predominantly accounted for by either changes in population shares (due to households moving across subgroups in such a way as to reduce the between-groups component) or by changes in relative sub-group mean incomes. In fact, in all those cases it is the unexplained effect (due to changes in inequality within the sub-groups over the period), which accounts for most of (or more than) the total change.

This can be interpreted to mean that the decline in inequality (as measured by the bottom-sensitive E(0)) between 1987 and 1994 is not primarily due to any reallocation of people across regions, or from rural to urban areas, or across age or occupational groups. Neither is it due to reductions in the disparities between the mean incomes of subgroups in those partitions. For all of those partitions, the dynamic decomposition assigns most of the change to the unexplained or residual 'within-group' changes. An insight into what may lie behind these within-group changes comes from the education partition: here, the most important component of the decomposition was a reduction in disparities between mean incomes across education sub-groups. Unexplained changes go in the same direction, but fail to completely offset the increase in inequality coming from the population shares effect, so that the overall change (-8.7%) is in fact less than the mean income effect (-10.8%). This suggests that the reduction in this bottom-sensitive inequality measure - which is most likely to reflect the compression at the lower tail of the distribution to which we referred in Section 3 - is due to a reduction in the returns to higher levels of education, caused by a gain in the incomes of the unskilled. The (inequality-augmenting) changes in the educational structure of the population (terms **a** and **b**) were insufficient to fully offset the effect of the lower returns.

partition variables (e.g. education, age, etc.). Note also that for some of the partitions, the four effects do not add up to the total change. That is due to the fact that the decomposition is an approximation, which is usually - but not always - very close to the exact decomposition. See Mookherjee and Shorrocks (1982).

So, does the Stolper-Samuelson Theorem apply to Chile after all, as Meller (1996) suggested? What does this new evidence imply with regard to the secular increases in the gap between the earnings of skilled and unskilled workers detected by Robbins and others in the 1970s and 1980s? While resolving this debate lies beyond the scope of this paper, we offer a simple hypothesis, drawing on the evidence just presented, as well as on previous research by Robbins (1996) and Wood (1997), that reconciles the evidence of increasing inequality until 1987-1990 and the broad stability of the mean-normalized distribution since then. This is, first, that insofar as the increase in earnings dispersion during the periods of active reform (1974 to the late eighties) was due to trade liberalization at all - as opposed to the deregulation in the labor market, for instance - this was due to its encouragement of what Wood (1997) calls 'skill-biased technical progress'. As Chilean firms opened up to competition with the outside world, and tariffs on capital goods were lowered, technological change took place which raised the demand for skilled workers. This effect, the data suggests, would have more than offset any static Stolper-Samuelson effect that might have been favoring unskilled labor.[28] Further evidence supporting this interpretation is supplied by Robbins (1995), who investigates the pattern of changes in employment in Chile over the 1957-1992 period, and finds that most movement took place within industries, towards more skilled employment (as one would expect from economy-wide skill-biased technological progress) rather than across industries (as one would expect from factor reallocation in response to price changes due to trade liberalization).

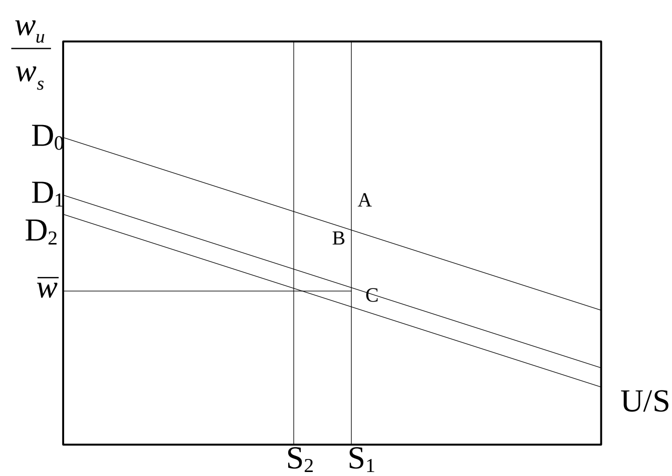
In terms of a diagram due to Leamer (1995), which was used by Wood (1997), such skill-biased technical progress caused the relative demand for unskilled workers to fall rapidly as trade liberalized, leading to the increased wage dispersion. As firms have caught up with

[28] Whether any such effect was indeed present or not remains unclear. The model predictions to which Robbins (1994) and others refer assume that the two types of labor are the only factors of production, and that all output is tradable. However, though the proportion has been declining, in 1996 83% of Chilean exports still consisted of minerals (largely copper), food and other agricultural raw materials. One could then reasonably suggest that land, rather than unskilled labor, might be the abundant factor in Chile. As for non-tradables, 55% of value-added (in 1980) was accounted for by services. See World Bank (1998). And as Wood himself writes, although the Hecksher-Ohlin results (including Stolper-Samuelson) are robust to the inclusion of more goods and countries, "the inclusion of non-traded goods and additional factors can yield contrary results in special cases" (Wood, 1997, p.36).

their international counterparts, and tariff levels have stabilized, the pace of skill-biased technical progress has slowed, however, and expansions in the supply of skilled labor, brought about by the large expansion in secondary and tertiary education in Chile over the last decade, now keep pace with demand.[29] See Figure 1 below.

So, whereas the economic reforms of the seventies and early eighties - of which greater trade openness was one aspect - led relative wages to fall from B to C, as relative demand for unskilled workers fell from D_0 to D_1, rather than rising from B to A, as the inherently static Stolper-Samuelson Theorem (for many goods) would have predicted; in the late eighties/ early nineties, slower rises in the relative demand for skilled workers are matched by increases in the supply of educated candidates. As demand shifts from D_1 to D_2, supply shifts from S_1 to S_2, keeping the relative wages roughly stable at $\overline{w}$.

Figure 1: Reaching Calm, after the Storm: a Dynamic Equilibrium in the Market for Skills, with Stable Inequality.



Note: With relative supply of unskilled to skilled workers given by S_1 a purely static trade liberalization process would have moved the economy from point B in the autarkic relative labor demand curve (dotted) to point A in D_0 , inducing a decline in wage dispersion. However, if lower tariffs on capital goods imports and stronger competition from abroad forces a skill-biased technical shift, the relative labor demand might shift to D_1 , so that the economy ends up at C, with a higher degree of wage inequality, as observed by Robbins (1994). However, after the initial catching up, technical progress slows to a shift from D_1 to D_2. If education shifts relative supply of skills from S_1 to S_2, relative wages might remain roughly stable.

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[29] Between 1987 and 1994, the population shares living in households headed by people with less than four years of schooling fell from 21% to 16%. Meanwhile, the share of heads with 12-15 years of schooling more than doubled from 11% to 25%.

As indicated in footnote 28, we do not really believe that this simple diagram represents a model capable of capturing the intricate dynamics of income distribution in Chile, which is a country with an important non-tradable sector, and exports which are at least as intensive in 'land' and natural resources as they are in unskilled labor. The links between international trade and personal income distribution are substantially more complex than that (as indicated by Spilimbergo et al, 1997)[30]. However, there might indeed be something to Wood's insight that what lay behind the secular increase in inequality before 1990 was skill-biased technical progress, accelerated by the competitive pressures of trade liberalization. This is supported by the employment evidence of Robins (1995).

It is also compatible with our findings of a roughly stable distribution of income in 1987-1994. Over this period, tariffs and other forms of protection had stabilized at their new, post-reform levels. The pace of increase in the demand for skills ought to have been reduced, just as educational spending increased, driving up the rate of increase in the relative supply of skilled workers. By itself, this would not only lead to a roughly stable distribution of incomes, as we observed. It would also lead to the observed declines in poverty, to the compression of incomes at the lower tail, and to the results described in Table 13: as people move from lower within-group inequality subgroups (with less education) to higher within-group inequality subgroups (with more education), the population share components of the Mookherjee and Shorrocks decomposition are positive. But this movement has a compressing effect on the ratio of mean incomes across education subgroups (with population shares constant). For a measure very sensitive to the bottom of the distribution, this has led to a decline in measured inequality. Overall, however, as we

[30] In preparing their aggregate results, Spilimbergo et al (1997) calculate an index of 'relative abundance' for various factors, given by the log of the ratio of the country's endowment of the factor to the world's average (*weighted by population and degree of openness*). According to their figures, Chile went from relative scarcity in both skill and land (-0.03 and -0.07, respectively) in 1966, to relative abundance (0.43 and 0.35) in 1990. In that year, in fact, these ratios would suggest that Chile had a comparative advantage in skilled-intensive sectors. These trends might support Wood's (1997) other hypothesis - that the entry of China and other large unskilled exporters changed the global patterns of comparative advantage. They also highlight the dangers of basing 'storm-in-a-tea-cup' academic debates on presumptions of factor endowments which are not supported by the data.

know from the Lorenz dominance analysis and from the behavior of the Gini coefficient and the Theil index, inequality was broadly stable - as represented by a stylized stable $\overline{w}$.

6. Conclusions

In presenting an overview of poverty, inequality and welfare trends in Chile during the last decade, this paper has sought to establish a number of empirical conclusions which can be drawn with confidence from a careful treatment of the data. these can be summarized as follows. Chilean inequality is high by international standards, and remained largely unchanged between 1987 and 1994. Scalar inequality measures - such as the Gini coefficient and the Theil index - varied little over the period, and there was no statistically significant Lorenz dominance between any of the years in the sample.

Within this broad picture of stability in inequality, however, there is some evidence that the shape of the density function may have altered slightly, with a compression at the lower tail (reducing bottom-sensitive inequality measures such as the mean log deviation), and an increase in dispersion at the upper tail (leading to rises in top-sensitive measures such as the coefficient of variation).

Economic growth has had a substantial and clearly beneficial impact, helping shift the distribution function to the right. Welfare, as measured by any reasonable social welfare function, was unambiguously higher in both 1992 and 1994 than it had been in either 1987 or 1990. Poverty was similarly incontrovertibly lower. The bulk of these welfare improvements took place between 1987 and 1992, however, and the slowdown in growth and increase in unemployment after 1992 were associated with losses to the poorest eight or nine percent of the population. While other social groups - including many of the poor - continued to gain, those losses at the very bottom of the distribution were associated with an increase in indigence in those two years.

Over the whole period, however, growth and other factors have led to a remarkable decline in poverty, however it is measured, and for a wide range of plausible poverty lines. For some of these lines, headcounts and poverty deficits were halved from 1987 to 1994.

While we have found that some assumptions underlying previous studies contributed to underestimating both poverty and inequality, whether by arbitrarily reducing prices in calculating poverty lines for rural areas, taking households rather than individuals as the unit of analysis, or ignoring live-in domestic servants, another important adjustment made in this study goes in the opposite direction and suggests that previous reports may have overestimated poverty and inequality. We have found that both poverty and inequality measures for Chile are considerably reduced when household incomes are adjusted to take into account differences in needs between children and adults, and to take some account of economies of scale inherent in the sharing of fixed costs within the household.

Finally, a preliminary investigation into the factors that may explain the structure of and the changes in inequality in Chile suggests that education is by far the most important candidate variable. Not only do differences between groups partitioned by educational attainment explain a much greater share of overall inequality than for any other household attribute, but also it is changes in returns to education which appear to lie behind the reduction in bottom-sensitive inequality measures over the period.

The importance of increases in education and the stability of overall inequality, taken together, also suggest that Chile may have reached a dynamic equilibrium between rising demand for and supply of skills. This would explain the recent reversal (or flattening) of the upward trend in both earnings and personal income dispersion that had previously been identified by various authors, for periods prior to 1990. With the profound political and economic changes of the 1970s and 1980s behind it, and quite regardless of its trade patterns, Chile’s income distribution is, for the moment, calm.

Appendix 1: The Data Set.

The CASEN sampling methodology can be described as multi-stage random sampling with geographical stratification and clustering. The country was first divided into strata comprising the rural and urban sectors of each of the 13 regions.[31] The rural sectors were final level strata. The urban sectors were further subdivided into three categories of towns, according to population: towns between 2,000 and 9,999 inhabitants; towns between 10,000 and 39,999 inhabitants; and towns with 40,000 or more inhabitants. All of the latter were sampled (i.e. they were final level strata). For other towns, there was a level of clustering in the selection of towns for sampling. At this level, with selected small towns, all large towns and the rural sectors, a first stage samples primary units ('zonas de empadronamiento'), with probabilities proportional to population. A second stage samples households. This process is described in more detail in Annex III to MIDEPLAN (1992). The sample sizes for our analysis were as follows: the 1987 sample included 23,403 households; the 1990 sample consisted of 26,248 households; the 1992 sample numbered 36,587 households; and the 1994 sample covered 45,993 households.[32]

Once each survey was completed, the data was entrusted to CEPAL (The United Nations Economic Commission for Latin America and the Caribbean), which conducted two types of adjustments to the raw figures. The first type comprises corrections for non-response, which are made in three instances: when people who declare themselves employed report no income from their main occupation; when people who state that they receive an occupational or widow's pension do not report a value for this benefit; or when owner-occupiers of their domiciles do not report a value for imputed rent. In all three cases, missing income values are replaced by the average value of the specific income variable in the household group to which the household belongs, where the group is defined by a partition according to a number of variables, including region, gender of head, educational attainment of head, occupational sector and category. See Appendix 1 to CEPAL (1995) for details.

The second type of adjustment seeks to correct for under- (or over-) reporting of different income categories, a common problem with household income surveys everywhere. For this purpose, CEPAL uses as the reference point for aggregate income flows the Household Incomes and Expenditures Account of the National Accounts System (SCN) of the Central Bank of Chile. First, a careful process is undertaken to convert the information in the original Central Bank accounts to the income concepts surveyed by CASEN. Once that is achieved, totals by specific income category are compared for CASEN (with recourse to the appropriate expansion weights) and the National Accounts. Finally, the proportional differences for each income category between the two sources are imputed uniformly to each income recipient in CASEN, with two notable exceptions: the adjustment in capital incomes is applied only to the top quintile (of households), proportionally to the primary

[31] For this purpose, an urban area was any grouping of dwellings with a population in excess of 2,000.

[32] These sample sizes are slightly larger than those reported in the official MIDEPLAN records of the surveys, reflecting our treatment of live-in domestic servants as separate households. See below.

incomes (ingresos autonomos) of all recipients there; and incomes from entitlement transfers and gifts are not adjusted.[33] The underlying assumption justifying this procedure is that mis-reporting differs fundamentally across income categories, rather than income levels.[34] In fact, the imputation would be strictly correct only if the income elasticity of mis-reporting within each income category was unitary. The only exception to this assumption, as noted, was in the treatment of capital incomes, which were imputed proportionally, but exclusively within the richest 20% of households. A detailed account of CEPAL's adjustment methodology, complete with the numbers used in each of the four years, is available in CEPAL (1995).

We have also made two adjustments to the data set, after it was processed by CEPAL. The income variable from the CASEN records on which our analysis is primarily based is total adjusted household income (YTOTHAJ), which includes all primary incomes, monetary transfers[35] and gifts, as well as imputed rent, after the CEPAL adjustments.[36] It was from this variable that we constructed both of the income concepts listed above - household income per capita and household income per equivalent adult - by appropriate choice of denominator. Our first adjustment was in the treatment of live-in domestic servants. It is unclear how previous studies treated them, since YTOTHAJ is defined to exclude their

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[33] It is suggested that the main reason for not adjusting these is that under-reporting of transfers consists mostly of complete omissions of benefits by some households, rather than proportional under-reporting of values by all recipients. There being no way to identify the omitters, no adjustment was found which would have improved the picture obtained from the survey.

[34] It may be interesting to note that the proportional adjustments did vary substantially across income categories. In fact, imputed rents were consistently found to have been over-reported, and were adjusted downwards in every survey.

[35] We believe that the questionnaire coverage of monetary transfers from the State, whether at the federal or municipal level, is exhaustive. Questions are asked and amounts are registered for the following benefits: Asignaciones Familiares, PASIS, SUF, Subsidio al Consumo de Agua Potable and Subsidio de Cesantia. See MIDEPLAN (1996) for a description of each of these entitlement benefits. Other benefits, associated with formal employment, such as jubilaciones, pensiones de invalidez and montepios are also included, although they are aggregated as part of primary - rather than secondary - incomes.

[36] We would have liked to use an even broader income concept, which took into account the values of the transfers in-kind which the government makes to many households, through programs in the areas of education, health and housing. The monitoring of these expenditures is in fact one of the objectives of the CASEN, and an exercise of valuation of these benefits is carried out by MIDEPLAN, relying on answers to survey questions about the usage of services and on cost data provided by the relevant ministries. Although the methodology for these valuations is discussed in MIDEPLAN (1993), and a tabulation of the imputations is available for quintiles of households in MIDEPLAN (1994), the disaggregated imputations at the household level - which would have been necessary for our analysis - are not made available with the other CASEN variables. This is due to alleged problems of methodology and reliability. The results are therefore only available at the quintile level, for distributions of households, and are of very limited comparability with the distributions used in this study. One should bear in mind that the - often substantial - value of these services, many of which are targeted, is omitted from the income data, when interpreting the distributional results that follow, or attempting to draw any conclusions about the social policies followed during the period.

incomes. Household-based studies are likely to have unwittingly excluded them from the sample altogether, by simply imputing YTOTHAJ to the household. For this Report, in households with live-in domestic servants, all other members received YTOTHAJ divided by the appropriate denominator (their number in the per capita income case, or the equivalence scale defined over them), while the servants were treated as a separate household, whose income was the sum of total adjusted individual incomes (YTOTAJ) over them.

The second adjustment was to exclude from the analysis the three richest households in the 1994 sample. This decision was carefully considered, and was based on the impression that these households reported sufficiently disproportionate incomes to be regarded as genuine outliers.[37] This impression was reinforced by the fact that two of these households were identical in every respect, having clearly been double-sampled[38], and by the position of the three households as outliers in a plot of the Pareto distribution of the top 1% of the sample.

[37] The value of E(2) - see below - when the outliers are included, is 6.58 in 1994. Its range over the other years is from 1.39 to 1.74.

[38] The practice of imputing all values from one household (randomly selected within the cluster) to another, when the latter has failed to respond, is often adopted as a way of maintaining representativeness within a small cluster. When the ‘doubled’ household are the Rockefellers, the practice requires revision.

Appendix 2: Inequality and Welfare Analysis for Per Capita Incomes.

For comparability purposes, this Appendix replicates some of the analysis described in Section 3 for household incomes per equivalent adult, using instead the distributions of household income per capita. This is not only intended to show that the main results of the paper are remarkably robust to the choice of unit, but also to enable comparisons with other studies found in the earlier Chilean literature.

Table A2.1 below is analogous to Table 3, and reports mean and median incomes, as well as the same four inequality measures, for the per capita distributions. Two changes from Table 3 are noteworthy. First, there is a decrease in the absolute values of the mean and median incomes for each year, which follows from the fact that there are many more households for which the denominators are reduced by the application of the equivalence scale used in Section 3, than there are households for which they are increased.[39] Second, all inequality measures are higher for this distribution than for the distribution of incomes per equivalent adult. This is in keeping with international experience, where it has been repeatedly found that the per capita income distributions generate upper-bound values for inequality, when compared to other assumptions about differences in needs and economies of scale within the household (see Coulter et al, 1992, and Ferreira and Litchfield, 1996). The reason is essentially that large households, or those with many small children, are re-ranked upwards from the per capita to the equivalised distribution, with the usual impact of reducing overall disparities.

Table A2.1: Descriptive Statistics: Household Incomes per Capita				
	1987	**1990**	**1992**	**1994**
Mean Income	55,367	63,293	75,371	78,281
Median	29,148	34,153	40,378	43,277
Gini	0.5603	0.5563	0.5534	0.5454
E(0)	0.5611	0.5495	0.5287	0.5212
E(1)	0.6349	0.6509	0.6551	0.6194
E(2)	1.3903	1.7447	1.6680	1.7121

Despite those level changes, there are no modifications to the perceived trends in inequality: the Gini still suggests a slight downward trend, but is basically stable. The Theil index, too, is trendless. The mean log deviation continues to indicate a reduction in bottom-sensitive inequality, and the E(2) continues to point to a gradual increase in top-sensitive inequality. This picture is broadly confirmed by the mean incomes and shares of each decile in the distribution of household income per capita, as revealed by Tables A2.2 and A2.3 below, respectively.

[39] Only single-person households would have an increased denominator, and would have lower entries in the distribution of incomes per equivalent adult than in the per capita distribution.

Table A2.2: Decile Mean Incomes: Household Incomes per Capita				
	1987	**1990**	**1992**	**1994**
Decile 1	6676.23	7661.76	10167.84	9990.29
Decile 2	12132.12	14324.69	17705.69	18262.61
Decile 3	16318.19	19167.72	23584.72	24338.92
Decile 4	20879.13	24503.26	29437.37	30984.11
Decile 5	26126.02	30633.76	36392.34	38848.29
Decile 6	32878.11	38440.16	45607.72	48727.70
Decile 7	42401.62	48828.75	58039.65	61619.54
Decile 8	57411.76	64844.24	76712.96	82953.61
Decile 9	87997.74	98060.98	112840.20	124861.60
Decile 10	250834.80	286449.80	343171.30	342230.30
Top Percentile	694357.20	860933.50	1078200.00	1026185.00

Table A2.3: Decile Income Shares: Household Incomes per Capita				
	1987	**1990**	**1992**	**1994**
Decile 1	1.21	1.21	1.35	1.28
Decile 2	2.19	2.26	2.35	2.33
Decile 3	2.95	3.03	3.13	3.11
Decile 4	3.77	3.87	3.91	3.96
Decile 5	4.72	4.84	4.83	4.96
Decile 6	5.94	6.07	6.05	6.22
Decile 7	7.66	7.71	7.70	7.91
Decile 8	10.37	10.25	10.18	10.60
Decile 9	15.89	15.49	14.97	15.95
Decile 10	45.30	45.26	45.63	43.66
Top Percentile	12.54	13.60	14.31	12.94

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