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Inequality in the distribution of personal income in the world: How it is changing and why

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Abstract. The variance in the logarithms of per capita GDP in purchasing-power-parity prices increased in the world from 1960 to 1968 and decreased since the mid 1970s. In the later period the convergence in intercountry incomes more than offset any increase in within country inequality. Approximately two-thirds of this measure of world inequality is intercountry, three-tenths interhousehold within country inequality, and one-twentieth between gender differences in education. If China is excluded from the world sample, the decline in world inequality after 1975 is not evident. Measuring confidently trends in household and gender inequality will require much improved data.

JEL classification: D31, F02, J16

Key words: World inequality, household inequality, gender inequality

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1. Introduction

Two empirical regularities in the distribution of income have recently gained the limelight in economics. The first is the tendency for income per capita – across countries, regions, states – to converge over time toward a steady state growth path, where convergence is associated with the (negative) estimated effect of initial income level on the subsequent growth rate, conditional (or unconditional) on inputs to growth – human capital, physical capital, research and development, government activities, and social and political conditions (Barro and Sala-i-Martin 1992, 1995). Since the Second World War there has been such an unconditional convergence across high-income countries and this tendency is also evident across subregions of the United States, Japan and Europe over relatively stable historical periods. This approach has been extended to a global scale, where institutional and technological possibilities differ more across countries, and the deterministic models are accordingly respecified to deal with inputs, stochastic growth, and country heterogeneity. The evidence for convergence is then more ambiguous (Dowrick and Nguyen 1989; Maddison 1989; Levine and Renelt 1992; Quah 1993; Durlauf and Johnson 1995; Lee et al. 1996; Williamson 1996).

The second empirical regularity is the increase in inequality in the distribution of personal income in many high income countries after 1980, which is particularly pronounced in the United Kingdom and the United States (Murphy and Welch 1992; Karoly 1993; Burkhauser et al. 1996; Gottschalk and Smeeding 1997 a). This growth in inequality is associated with increased wage differentials by skill, measured by schooling, occupation, and labor market experience, but not necessarily by gender. The growing importance of international trade is ascribed a role in the intercountry diffusion of this change in wage structures, but not all economic studies confirm an important role of international trade compared to the residual skill-biased technical change (e.g., Burtless 1995; Blau and Khan 1996).

The first intercountry convergence implies decreasing inequality across a subset of relatively rich countries, whereas the second empirical regularity reflects increasing inequality within the same countries. One objective of this paper is to bring these two pieces of evidence together to describe how inequality has evolved across this increasingly integrated group of advanced economies. Moreover, there appear to be sufficient data to extend tentatively the analysis to the less advanced economies and ascertain whether global forces are at work in these countries, as well, promoting intercountry convergence and increasing intracountry inequality. It is also common in both of these literatures to treat regional subeconomies, countries, or administrative units within countries as equivalent observations in growth (or inequality) regressions. For my purposes, however, it is more reasonable to weight countries by their populations. This natural shift to population weighted comparisons has obvious implications for the importance assigned to the growth of, and inequality within, the largest countries, such as China and India.[1]

Approximating the personal distribution of income or welfare in the world requires four types of data for all countries: the population size, the income level, the interhousehold distribution of income, and the intrahousehold distribution of welfare. It is not difficult to understand, therefore, why there are relatively few descriptions of the global inequality. The quality and comprehensiveness of information tends to diminish as one descends this shaky empirical ladder. Although scattered references are found to the widening gap in income between the rich and poor countries (e.g., UNDP 1992; Quah 1993; Pritchett 1996, 1997), empirical studies are rare.[2] The analysis of Berry et al. (1983, 1991) for the period 1950 to 1977 provides a firm starting point. Twelve to seventeen more years of intercountry data and an increased range of evidence along other dimensions may justify revisiting these issues. This paper is designed, consequently, to initiate discussion and interpretation of the available evidence, to identify data gaps and close them, where necessary, with one possible set of working assumptions. Kuznets (1955, 1963), among others, proceeded several decades ago to initiate analyses of the personal income distribution within nation states. This paper starts putting those national estimates together to see if the global consequences of economic and demographic growth for the distribution of income among the world's people can be quantified and related to the mechanisms of development.

The balance of the paper is organized as follows. Section 2 outlines my approach for decomposing the log variance of personal incomes by three levels of aggregation (country, household, gender), and contrasts this to other measures of inequality, other income units, and welfare. Section 3 describes the intercountry inequality in income distribution. Section 4 examines how these trends have been shaped by regional patterns. Section 5 reports how intracountry inequality is estimated, with all of its uncertainty, and incorporates this component into estimates of world inequality. In Sect. 6 the gender inequality within the household is assessed and factored into the total. Section 7 concludes with ideas for extending and improving the data and methods and on the possible connections between inequality and growth.

2. Data and measurement

Refinements in demographic estimates in the last several decades have created a consensus as to national population figures. Although the size and age composition of some national populations may not be known with great precision, indirect methods for estimating vital rates based on analyses of age compositions have narrowed demographic uncertainties substantially (United Nations 1967), while coordinated series of national household surveys have improved our knowledge of fertility and child mortality in low-income countries, and these vital rates account for much of the variation in population growth. I have used the population estimates in Summers and Heston (1991) Penn World Tables (Mark 5.5) which generally replicate World Bank Tables and are usually similar to those published by the United Nations Population Division.[3]

There are two widely consulted measures of national income that differ in terms of how local currency constant price national income accounts are compared across countries.[4] The traditional approach was to use foreign exchange (FX) rates between each country and a numeraire currency, such as the US dollar, to arrive at equivalent foreign trade purchasing power.

This approach has three limitations: First, not all goods are traded (e.g., housing and many services), second, foreign exchange markets are often regulated; and third volatile capital movements add swings in foreign exchange rates that may not well approximate the personal consumption opportunities provided by local income. In the last decade local currency price indices have been developed to provide an alternative approach to link currencies in terms of purchasing power parity (PPP) for a common bundle of goods (Kravis et al. 1982; Summers and Heston 1991). FX rates might mirror PPP rates between currencies, but traded goods as a share of income have increased over time, and differ across countries, whereas capital flows and expectations about macroeconomic policies can interject discontinuities in FX rates as will be seen in East and Southeast Asian incomes after 1997. Smaller countries may be linked in this process to the fate of the currency of their major trading partner, with repercussions of changes in FX rates among dominant currency countries diffusing to the periphery, such as from France to Francophone Africa, or from the United States to Latin America and recently Thailand. Both the FX and PPP estimates of per capita income are reported below. There are still methodological issues concerning the concepts and classifications applied in generating the national price indexes for the core benchmark countries, such as how to treat the quality of untraded services, as well as the approach used to extrapolate price indexes from this benchmark sample to the 120 countries examined here (Maddison 1989; Heston 1994; Bernard and Jones 1996). Nonetheless, given my goal to describe trends in the distribution of personal welfare that are due to income, the general concept of purchasing power parity is the appropriate one for this study.

The universe examined here is limited by the availability of national income accounts. Estimates are available for Gross Domestic Product (GDP) for 120 countries from 1960 to 1989 from the Penn World Tables (Mark 5.5). Many countries outside of the OECD and Latin America do not have estimates of national income before 1960, and the panel can be extended beyond 1989 for only a few countries. These 120 countries contain 93% of the world's population in 1960 and 92% by 1989 (see Appendix Table A-1 for a listing).

The third set of data are national estimates of the size distribution of household income. These data depend on the concept of income and the definition of the income unit, neither of which is widely standardized. Nor is there a consensus on how to translate household income into an indicator of average personal welfare for its members, as the composition of households will vary across countries and change over time within them, in response to socioeconomic conditions that include income (Schultz 1997). Income distribution data compiled by Deininger and Squire (1996) are analyzed if a country provides at least two national representative samples since 1950. The 56 countries included in the working sample include 83% of the world population in 1960 (see Table A-1). Estimates of the log variance, Gini concentration ratio, and Theil mean log deviation are estimated on the basis of the cumulative share of income received by the first four quintiles of the income units.[5]

The fourth level of data relates to intrahousehold inequality, and it is not currently collected, allowing me more scope for imagination. The distribution of resources within the household has only recently begun to receive systematic study by economists (McElroy and Horney 1981; Schultz 1990; Thomas 1990, 1994; Chiappori 1992; Bourguignon et al. 1993; Hayashi 1993; Browning et al. 1994). Nash-bargained or Pareto-efficient sharing rules have been used to interpret variation in the intrahousehold allocation of resources among members. In this context it has been hypothesized that the earnings opportunities of men and women outside of the household may affect the resources they control within the household, by changing the member's "threat points", even when partners do not actually enter these labor markets that are external to the family. With schooling being the most influential explanatory variable for wages of men and women, I focus on the gender gap in schooling as a proximate determinant of the gender gap in personal income or welfare (Schultz 1993). Although education may be arguably the most important measurable aspect of gender inequality, it should be supplemented when reliable data are widely available on gender differences in health, wages, and consumption, and their correlation between spouses and within households. Other aspects of intrahousehold inequality might focus between generations of adults in extended families, or between parents and children, but I know of no data assessing intergenerational inequalities across a sample of countries. However, in families where women are better educated, children do tend to be healthier and better educated, controlling for the family's income per capita, while fertility and population growth tend to be lower (Schultz 1993; Thomas 1994).

Educational attainments by sex have recently been estimated by country in several studies. Schultz (1987) analyzed the determinants of expected years of schooling by sex, based on period-specific enrollment rates from 1960 to 1980 summed over levels of schooling. Barro and Lee (1994) estimated for every five years from 1960 to 1985 the mean years of educational attainment for men and women over age 25 for 129 countries from UNESCO tabulations of educational attainment by age and sex. Dubey and King (1994) estimate educational stocks by sex and age for 85 countries from 1960 to 1987 using cohort enrollment models. This paper relies primarily on the Barro and Lee estimates, which include the largest number of countries. An expected current enrollment level for women and men is also estimated from UNESCO data, and it is used later as an alternative basis for assessing the effects of gender inequality on economic growth.

Measurement issues

Inequality is measured in many ways and some have more attractive features than others in terms of decomposing aggregate inequality into between and within subgroup components and satisfying reasonable economic restrictions (Kuznets 1963; Sen 1973; Shorrocks 1980; Cowell 1995; Morduch and Sicular 1996). Because the frequency distribution of households by log income is often approximately normal, the variance of the logs of income is a parsimonious description of the distribution of income that is unit free and is used in this paper for comparative purposes. To assess whether trends over time in inequality depend on the measure consulted, I report also Theil's (1976) second measure of entropy or the mean log deviation, which weights subgroups by their populations, and the Gini concentration ratio that averages all individual differences and is visualized in terms of the Lorenz diagram. The axiomatic method for selecting an index of inequality that satisfies reasonable economic properties and is subgroup additively decomposable eliminates many traditional measures including the log variance and Gini, and leaves only three candidates: the two based on entropy (Theil 1967) and the squared coefficient of variation (Bourguignon 1979; Shorrocks 1980). A heuristic log variance decomposition is described below, although it is additively decomposable only when inequalities are orthogonal across levels, as in the randomized treatment model of Fisher (1930), and the Theil mean log deviation decomposition is shown between and within countries in Appendix Table A-3.[6]

Let Y_{ijk} be the natural logarithm of an adult's income in the ith country ($i=1,2,\ldots,c$), in the jth household ($j=1,2,\ldots,h_c$), of the kth gender ($k=1,2$). The mean log income is defined as follows:

$$\bar{Y} = \sum_{i=1}^{c}\sum_{j=1}^{h_c}\sum_{k=1}^{2} Y_{ijk} \Bigg/ \sum_{i=1}^{c}\sum_{j=1}^{h_c}\sum_{k=1}^{2} n_{ijk}\,,$$

and the variance of the logarithms of income, σ_y^2, is a unit-free measure of inequality:

$$\sigma_y^2 = \sum_{i=1}^{c}\sum_{j=1}^{h_c}\sum_{k=1}^{2} (Y_{ijk} - \bar{Y})^2 \Bigg/ \sum_{i=1}^{c}\sum_{j=1}^{h_c}\sum_{k=1}^{2} n_{ijk}\,,$$

where n_{ijk} is the number of adults with Y_{ijk} income. A linear model is assumed with country-, household- and gender-income effects that operate independently on the logarithm of personal income, where it is commonly assumed that the log income variable is distributed normally.

Interactions among the three levels of classification are neglected. This measure of inequality can then be decomposed into (1) international differences *between country's* means and the world mean, squared and weighted by the country's population, plus (2) the *within country* log variance across household, weighted by population, and (3) a *within household* log income variance, weighted by population:

$$\begin{aligned}\sigma_y^2 = {} & \left[\sum_{i=1}^{c} n_i\,(\bar{Y}_{i..} - \bar{Y})^2\right]\left(\sum_{i=1}^{c} n_i\right)^{-1} + \left[\sum_{i=1}^{c}\sum_{j=1}^{h_c} n_{ij}\,(\bar{Y}_{ij.} - \bar{Y}_{i..})^2\right] \\ & \times \left(\sum_{i=1}^{c}\sum_{j=1}^{h_c} n_{ij}\right)^{-1} + \left[\sum_{i}\sum_{j}\sum_{k=1}^{2} n_{ijk}\,(Y_{ijk} - \bar{Y}_{ij.})^2\right] \\ & \times \left(\sum_{i=1}^{c}\sum_{j=1}^{h_c}\sum_{k=1}^{2} n_{ijk}\right)^{-1}. \end{aligned} \tag{1}$$

The first of the three components of the variance is calculated from national income and population data.[7] The second variance component requires estimates of the log variance of interhousehold income inequality within countries, where income in the household would be ideally measured on a per capita or per adult basis. The third component of variance allows for intrahousehold inequality as subsequently approximated by human capital differences by gender, one possible indicator of individual productivity and bargaining power.[8] The difference between the log income at five quantiles and the log of mean income for the distribution of PPP and FX incomes is reported in Appendix Table A-2 to provide information about which quantiles in the distribution are changing.[9]

A second indicator of income inequality is computed from the intercountry data for comparison purpose, although no decomposition is reported. The Gini concentration ratio (G) is defined as the sum of the absolute value of the differences in income, y, between all possible pairs of households, divided by the product of twice the mean income (m) and the total number of households (n) squared:

$$G = [2\,mn^2]^{-1} \sum_{i=1}^{n} \sum_{j=1}^{n} |y_i - y_j| f\,(y_i) f\,(y_j)\ , \tag{2}$$

where the subscripts i and j run across all n households, and $f(y_i)$ is the number of households (or adult population) with income y_i.

Finally, the second Theil (1967) entropy index of inequality, often called the mean logarithmic deviation, is defined as the sum of the log of the world population mean income relative to the country mean incomes, weighted by the country's population share:

$$T_2 = (1/n) \sum_{i}^{n} log_e\,(m/y_i)\,, \tag{3}$$

which can be decomposed across c countries and j quantiles of households within countries as follows:

$$T_2 = \sum_{i}^{c} p_i\ log_e\,(p_i/s_i) + \sum_{i}^{c} p_i\ \sum_{j}^{h_c} (p_j/p_i)\ log_e\,((p_j/p_i)/(s_j/s_i))\ , \tag{4}$$

where p_i and p_j refer to the shares of world's population in the ith country or jth quantile of households in that country, and s_i and s_j refer to the shares of world's income received by the ith country or share of the country's income received by the jth quantile of households. The first term on the right is the intercountry component of income inequality and the second term is the interhousehold component of income inequality within each of the countries, weighted by the country's relative size of population.

The variance in the logs of potential earnings between women and men can be related to the gender difference in completed education. The structure of wages of men and women workers has been summarized in many countries by fitting them to a log-linear wage function, following Mincer (1974):

$$Y_k = a_k + r_k E_k \, , \quad k = 1, 2 \, , \tag{5}$$

where Y is the logarithm of the opportunity wage (or earnings of the individual given comparable potential labor supply), E is years of schooling, a controls for the wage effect of other observable productive factors, where the index k denotes gender, and the individual subscripts have been suppressed for simplicity. The private wage return to years of schooling, r_k, is merely the percentage increase in wages associated with an additional completed year of education, and when estimated for women and men these returns are of roughly similar magnitudes for the same levels of schooling, or women's returns are slightly higher than men's (Schultz 1993). It is assumed here that the parameters r and a are equal for men and women and that the wage effects of education and other factors do not interact. If husbands and wives were perfectly positively sorted by schooling, so that the man with the most education marries the woman with the most education, within the relevant age group, and so on down the distribution of education, the Pearson correlation of spouses education would be perfect and $\rho = 1.0$. But empirical estimates of this correlation generally fall in the range of 0.4 to 0.7 (Mare 1991; Kremer 1997).[10]

The variance in log potential earnings of men and women can then be expressed as a product of the squared average gender difference in schooling and the wage return on schooling squared:

$$V\,(Y_k) = (1/2\,\rho)^2\, r^2\,(E_k - E_k)^2 \, . \tag{6}$$

Lacking estimates of ρ and r for virtually all countries, the working assumption is made that $\rho = 0.5$ and $r = 0.15$ for all countries. The resulting rough indicator of the contribution of gender differences in education to the log variance in gender log wage opportunities is thus obtained:

$$V\,(Y_k) = \sim 0.0225\,(E_1 - E_2)^2 \, . \tag{7}$$

Clearly, this approximation is most crude and should be derived through analyses of individual survey data from each country. Even when satisfactory data is available for this purpose, many analytical issues remain to be resolved. The selection process determining who is married with spouse and who works for a wage must be jointly modeled in order to estimate unbiased wage opportunities for all persons. Single adult households would also contribute directly to household income inequality, and need to be included in the share of inequality due to gender differences in earnings potential. Including gender differences in health and capabilities would complicate further the measurement problem (Sen 1973). The above approximation is only offered as a starting point for much further conceptual and empirical refinement.

A few examples suggest the range and magnitude of this approximation of $V(Y_k)$ across countries and over-time within countries. In India in 1980 the average female and male adult schooling was 1.4 and 4.0 years, respectively, implying a log variance of gender earnings of 0.152, whereas in Indonesia in the same year women and men reported 2.2 and 3.9 years for a log variance of gender earnings of 0.056, roughly a third the level of India (data from Barro and Lee 1994). Men and women in Taiwan who were born between 1917 and 1921 and survived to 1967 had an average difference in schooling of 4.2 years, whereas those born between 1966 and 1970 who were surveyed in 1995 had a gender gap in schooling of 0.23 years. According to my approximation the direct contribution of gender differences in schooling to variance in logs of personal earnings potential in Taiwan would have declined from 0.388 to 0.001 in this fifty year period.[11]

3. Intercountry inequality in per capita incomes: Trends and population growth

Table 1 summarizes the distribution of world income based on the per capita Gross Domestic Product (GDP) estimates for 120 countries in a thirty-year period.[12] The dispersion of incomes within countries or households is initially ignored in the intercountry inequality measures, and everyone in a country is implicitly being attributed the average income for that country. The population weighted variance in log income per capita (in 1985 US dollars), based on the traditional foreign exchange (FX) rate equivalence of local currencies, increased 60% from 1.51 in 1960 to 2.42 in 1989, whereas, according to their currency's purchasing power parity (PPP), the thirty year increase in variance in log per capita income is only 7%, from 0.943 in 1960 to 1.01 in 1989. These series are plotted as the middle lines in Fig. 1, and are bracketed by the series in which the population weights of countries are held constant at their initial 1960 values (top line) and at their final 1989 values (bottom line). The Theil entropy index increases by nearly the same amount, 57% in FX income and 6% in PPP income. This parallelism is hardly surprising, for the entropy index is also based on deviations of log national incomes from log world income, and, for example, comparing PPP incomes, the log variance and the entropy index are correlated at 0.98 (Table 1). The Gini ratio, plotted in Fig. 2, increases more moderately by 13%, from 0.640 in 1960 to 0.725 in 1989 based on FX income, and advances only 1% in PPP income from 0.547 to 0.552.[13]

Trends in intercountry income inequality vary in the period studied, but within the same concept of income (FX/PPP), the three summary measures of inequality imply concurrent time series variations. Movements in the quantiles of the distribution of incomes in the world are reported in Appendix Table A-2. In terms of my preferred summary measure of inequality that lends itself to the later disaggregated decomposition, the PPP income log variance increases sharply in the first few years, 1960 to 1962, from 0.94 to 1.10, gradually rises further to its peak of 1.20 in 1968, returns to 1.19 in 1976, and thereafter declines until 1985 when it fell to 0.97 before stabilizing around 1.01. In sum, the log variance in PPP incomes rises by a fourth in the first decade, and then declines by a fifth in the final decade of my data.

Table 1. Inequality in intercountry income per capita: 1960–1989

Year	Variance of log income[a]		Theil entropy index[b]		Gini concentration ratio[c]	
	FX[d] (1)	PPP[e] (2)	FX (1)	PPP (2)	FX (1)	PPP (2)
1960	1.511	0.943	0.837	0.534	0.640	0.547
1961	1.633	1.050	0.886	0.574	0.648	0.599
1962	1.723	1.100	0.923	0.595	0.655	0.566
1963	1.702	1.075	0.907	0.594	0.651	0.563
1964	1.671	1.067	0.899	0.584	0.649	0.563
1965	1.635	1.065	0.889	0.583	0.648	0.566
1966	1.742	1.088	0.930	0.603	0.656	0.569
1967	1.820	1.131	0.969	0.618	0.662	0.573
1968	1.903	1.199	1.010	0.644	0.670	0.581
1969	1.868	1.151	0.995	0.626	0.667	0.575
1970	1.847	1.107	0.982	0.603	0.664	0.565
1971	1.890	1.111	1.002	0.606	0.668	0.566
1972	2.013	1.158	1.048	0.629	0.672	0.572
1973	2.055	1.187	1.056	0.641	0.669	0.574
1974	2.029	1.180	1.030	0.629	0.663	0.568
1975	2.023	1.136	1.032	0.607	0.666	0.561
1979	2.170	1.192	1.083	0.630	0.674	0.568
1977	2.151	1.166	1.076	0.619	0.674	0.566
1978	2.362	1.143	1.162	0.612	0.685	0.564
1979	2.341	1.149	1.155	0.614	0.684	0.564
1980	2.257	1.087	1.112	0.582	0.680	0.553
1981	2.279	1.071	1.127	0.578	0.682	0.553
1982	2.208	1.054	1.109	0.567	0.683	0.548
1983	2.123	1.029	1.050	0.559	0.689	0.546
1984	2.102	1.014	1.113	0.556	0.695	0.548
1985	2.002	0.970	1.086	0.535	0.694	0.540
1986	2.224	1.001	1.207	0.554	0.710	0.548
1987	2.427	1.019	1.304	0.562	0.722	0.551
1988	2.435	1.014	1.321	0.562	0.726	0.551
1989	2.419	1.011	1.311	0.563	0.725	0.552

[a] First term on right side of Eq. (1).
[b] Eq. (3).
[c] Eq. (2).
[d] Local currency real income (GDP) converted to US 1985 $ by foreign exchange rates (FX).
[e] Local currency real income (GDP) converted to US 1985 $ by purchasing power parieties (PPP).

One might imagine that these trends in inequality could be affected by the exceptional geographic distribution of population growth during this period, which reached its historic peak growth of 2.4% per year in 1960–1965, before falling to 1.7% by the end of this period. Although opinions vary widely, no satisfactory method has been developed to disentangle how this reduction in population growth facilitated economic growth in output per capita, and thus how it may have altered directly differences between countries in per capita incomes (National Academy of Sciences 1986). But three simple decompositions may capture some implications of the demographic transition for the world's intercountry income inequality. First, population relative

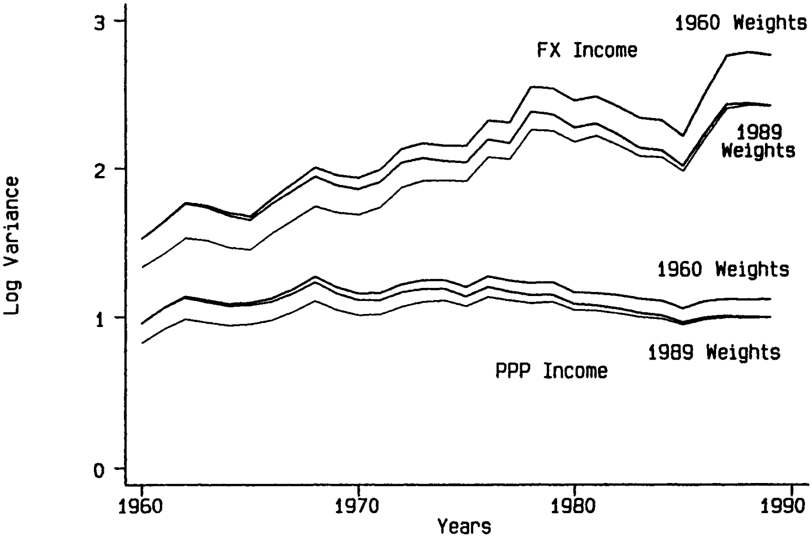

Fig. 1. The log variance of intercountry incomes per capita

weights of countries can be held constant, say at their initial or final year levels. Second, the rates of national population growth can be assumed to continue unabated at their 1960–1965 peak levels until 1989. And third, the changing age composition of populations that follows from the demographic transition can be used to refine our measures of national welfare.

In the first scenario, if the relative population weights of all countries are held constant at their 1960 levels, the log variance of intercountry PPP incomes as plotted in Fig. 1 would have been 13% higher in 1989 than with the actual changing weights, and 23% higher if FX incomes are examined. By holding constant the initial population weights, the weight is increased for the outlying high-income countries which in reality fell from one-third of the world's population to one-fourth in this 30-year period. In other words, the world's richest countries sustained below average population growth rates in this period, and they also experienced above world average rates of economic growth until 1975, and below average economic growth thereafter.

In the second simulation, the national rates of population growth recorded in 1960–1965 are assumed to have continued through 1989, whereas in reality these unprecedented rates of population growth decreased rapidly in Latin America and East Asia, and decreased slowly in South Asia, while they increased slightly in Africa on average, where child mortality fell faster than fertility. The resulting increase in the population weights of Latin America and East Asia is associated with an increase in the log variance in PPP incomes per capita compared with those reported in Table 1 based on current population weights, but the differences are only a few percent.

The third demographic-based simulation recognizes that children have lower consumption requirements than adults. Therefore, when the propor-

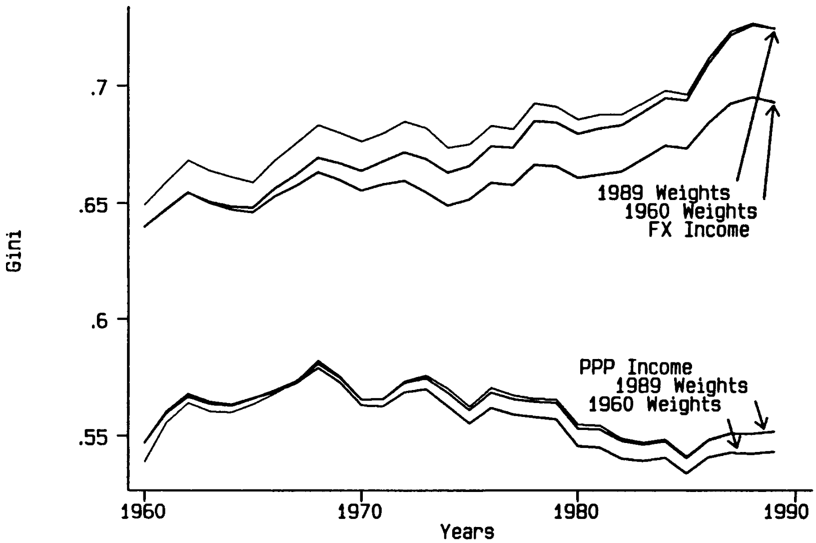

Fig. 2. The Gini concentration ratio of intercountry incomes per capita

tion of children in the population increases, as it did at the start of the demographic transition as child mortality declines, changes in per capita income understate the advance in welfare, whereas later as fertility declined and the proportion of children in the population decreases, changes in per capita income overstate the advance in welfare. To assess how important these changes in age composition are for measuring the level and trends in inequality, one can express national income in per adult units rather than per capita, although this approach undoubtedly understates adult equivalents but defines a maximum adjustment that might be defended to take account of the welfare effects of changing age compositions. According to this logic, the gains in per capita income would relatively overstate economic welfare advances in countries such as Taiwan and Korea where fertility fell by more than half after 1960, relative to India and Pakistan where fertility fell more slowly.

The log variance in PPP income *per adult* is 13% lower in 1960 than that of income *per capita*, since the proportion of children in the population is much higher in the lower income countries. By 1989 this measure of PPP income inequality per adult is 15% lower than per capita inequality. Thus, relying on a welfare indicator that focuses only on income per adult would imply that world log variance in PPP income increased more modestly than recorded earlier over the 30-year period, rising only 5.8% compared with the benchmark increase of 7.2% shown in Table 1 and Fig. 1. The demographic transition as it impacts on the relative weights of poor and rich countries reduced slightly measured world inequality in per capita income, and to the extent that adults have higher consumption requirements than children, the resulting decline in the child fraction of the world's population would have implied a lower level of inequality and a slower growth in inequality over time. All three simulations suggest that the changing population composition of the world was not a major factor behind the trends shown in Table 1 column 2, although they appear to have moderated any increase and amplified slightly the declines in PPP income inequality that began to emerge in the second half of the period.

4. Regional factors in intercountry income inequality

Table 2 reports for four years the mean and variance among countries in log GDP per capita based on the preferred purchasing power parity (PPP) methodology, first for the world and then for five subregions or groups of countries. Countries outside of the high income group (Eastern Europe and OECD) are divided into Latin America, South and West Asia (Bangladesh to Lebanon), East and South East Asia (China to Myanmar), and Africa. Figures 3 and 4 plot the annual mean and variance, respectively, for the five regional groupings, plus the consolidated low income country total, displaying both foreign exchange (FX) and PPP figures. The mean incomes illustrate the abrupt effect of foreign exchange crises in regions, such as Africa and Latin America, on the growth in incomes evaluated at foreign exchange rates, and the more smoothed path of PPP income. In Africa FX income declines sharply after 1980, whereas PPP income remains constant. In Latin America FX income dips after 1980 as the Mexican debt crisis ushers in a decade of stagnation in the region based on FX income, but modest growth continues based on PPP income. South Asia experiences more steady growth, with the exception of modest setbacks in 1966–1974 as the Indian subcontinent experienced agricultural reversals. East Asia also evidences the repercussions of China's famine of 1959–1961 and cultural revolution in 1966–1974. The high income country group is much less subject to swings in income measured on the basis of FX, although the second oil price shock and business cycle stopped FX income growth in 1980–1983, and individual countries experience periods when FX and PPP incomes deviate more widely.

Because of the greater homogeneity in income levels within a region than in the world, the intercountry variances tend to be substantially lower within the regions than across all countries in the world (cf. Theil 1967). Intercountry variances in log PPP income are increasing in Africa from 0.2 to 0.4, and in East Asia from 0.1 to 0.2, and decreasing within the high income countries from 0.49 to 0.25, as the recent convergence literature has stressed (Barro and Sala-i-Martin 1995). The intercountry convergence in incomes within the high income group halts after 1980, as measured by the variance in PPP log incomes, and starts to diverge in FX units, mostly because of the relative decline in FX income per capita in the USSR, Turkey, Yugoslavia, Czechoslovakia, and Greece, for example, as well as the increased relative FX income deviation of Japan. Foreign exchange market distortions as well as erratic economic policies could be responsible for a further deterioration in Eastern European fortunes after 1989, causing more divergence in FX incomes, and possibly even some divergence in PPP incomes in the high income group of countries (see also with entropy inequality in Table A-3). In Latin America the intercountry variance in log PPP income has been roughly constant at 0.15, while it has increased in

Table 2 Regional patterns in intercountry log variance of per capita PPP incomes

	Log income mean (1985 US $ per capita)	Log variance components			Population (billions)
		Intercountry	Intracountry	Total	
	(1)	(2)	(3)	(4)	(5)
1. World Total					
1960	5.93	0.943	0.473	1.416	2.812
1970	6.45	1.107	0.459	1.565	3.432
1980	7.41	1.087	0.437	1.524	4.132
1989	7.95	1.011	0.430	1.441	4.821
2. High income countries (OECD plus rest of Europe including Turkey)					
1960	7.07	0.491	0.461	0.952	0.909
1970	7.82	0.349	0.465	0.814	1.017
1980	8.83	0.243	0.428	0.671	1.110
1989	9.43	0.252	0.441	0.693	1.186
3. Africa (North and Sub-Saharan)					
1960	5.37	0.213	0.829	1.042	0.240
1970	5.85	0.268	0.799	1.067	0.307
1980	6.78	0.392	0.767	1.159	0.401
1989	7.01	0.415	0.740	1.155	0.521
4. Latin America (and Caribbean)					
1960	6.43	0.153	0.971	1.124	0.207
1970	6.98	0.151	0.952	1.103	0.273
1980	8.09	0.150	0.914	1.064	0.347
1989	8.37	0.147	0.894	1.041	0.418
5. South Asia (Bangladesh to Lebanon)					
1960	5.43	0.064	0.346	0.410	0.576
1970	5.85	0.124	0.335	0.459	0.733
1980	6.69	0.184	0.320	0.504	0.935
1989	7.36	0.095	0.302	0.397	1.153
6. East Asia (Korea to Myanmar)					
1960	5.11	0.067	0.325	0.392	0.887
1970	5.63	0.115	0.318	0.433	1.103
1980	6.75	0.156	0.304	0.460	1.339
1989	7.46	0.194	0.287	0.481	1.543
7. Low income countries (3+4+5+6)					
1960	5.38	0.249	0.465	0.714	1.910
1970	5.88	0.307	0.456	0.763	2.415
1980	6.89	0.383	0.440	0.823	3.022
1989	7.47	0.317	0.427	0.744	3.636

South Asia reaching a maximum in 1976 of 0.24, before declining to 0.10 in 1989. Combining the countries not in the high income group (or simply low income countries), one observes an increase in intercountry log variance in income from 0.25 in 1960 to 0.38 in 1980 before falling to 0.32 by 1989. The population weights associated with the regions are reported in the last column in Table 2.

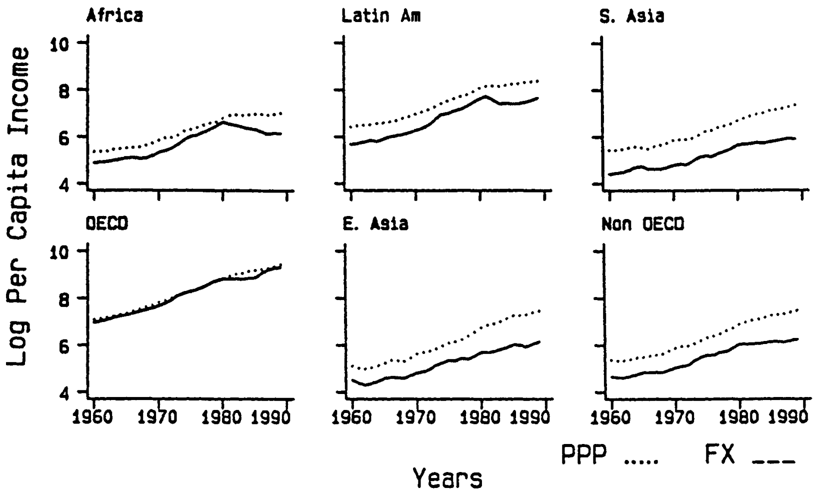

Fig. 3. Time trends in log per capita income by region

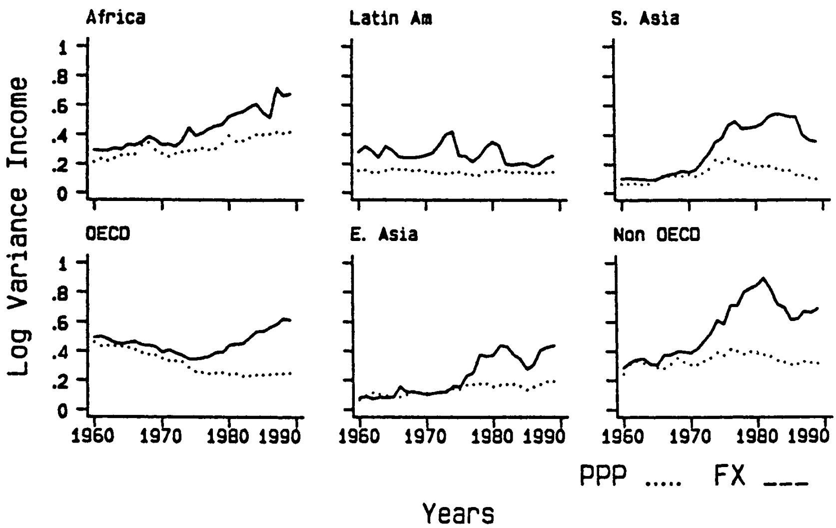

Fig. 4. Time trends in variance of log per capita income by region

China and India

China and India, the two largest populations whose incomes are substantially below the world's average income in 1960, make a major contribution to these summary measures of intercountry income inequality. Excluding China from the world sample reduces the log variance in PPP income per capita by 4% in 1960 but increases the log variance by 14% by 1989. This reflects the fact that China had a relatively low income in 1960 and grew more rapidly than the world average income after the mid 1970s. The log variance of PPP income per capita would therefore have been one-fifth higher in 1989 than in 1960, had China been excluded from the working sample, or stated in another way, the growth in Chinese income after the 1970s offset a marked increase in the world's inequality excluding China. Outside of China, income inequality peaks in 1968 and is relatively constant after 1976. Because China reduced its fertility sharply after 1970, whereas Indian fertility has fallen more gradually, the two countries begin to exhibit in this period quite different proportions of children. Expressed as income per adult, the exclusion of China again lowers the log variance in PPP income by 9% in 1960, and the time trends are similar as with per capita income, increasing by 22% to peak in 1976, and then declining gradually a few percent by 1989. Four-fifths of the decline from 1976 to 1989 in the log variance in income per adult in the entire world is accounted for by the inclusion of China in the sample. Whatever claims can be advanced for a reduction in world income inequality from 1974 to 1989 depend on the growth achieved by China in this period.

India has a smaller effect on the levels and trends in world inequality. Excluding India from the sample increases the log variance of PPP income per capita by 10% in 1960, by 6% in 1976, and by 8% in 1989, thereby reducing by 1.5 (1.3)% the increase in the log variance of income per capita (per adult) over the entire time period. Thus the inclusion of India decreases the level of world intercountry inequality, and augments slightly the growth in inequality over time, but does not alter markedly the overall trends or variations in inequality in the subperiods. India is a stabilizing force compared with China, whose volatility modifies world trends in different subperiods, from the famine following the "great leap forward" from 1959–1962, to the cultural revolution in 1966–1974, to the agricultural household responsibility reforms starting in 1979, and the subsequent rapid decentralized industrial expansion.

5. Intracountry income inequality

Personal income distribution estimates have been recently consolidated by Deininger and Squire (1996), in which they include 682 observations by country, year, income type, and form of recipient unit. All national observations that report income or total expenditures for households, or income for persons, on either a pretax (gross) or after-tax income basis are initially analyzed here. A further restriction is imposed that each included country provides at least two observations for them to contribute symmetrically to the information used for the pooled sample and the within-country estimates that include country fixed-effects. The maximum-sized working sample thus defined includes 509 observations from 56 countries that represent nearly four-fifths of the population in my 120 country sample.[14] Regressions were then estimated to account for the pooled year/country observations on the variance of the logs of income, the Gini concentration ratio, and the Theil entropy index, where Huber (1967) standard errors are reported to correct for heteroscedasticity across countries. The same variables were statistically significant in accounting for all three measures of inequal-

Table 3. Regressions for the log variance of intracountry incomes: Pooled and with country-fixed effects[a]

Explanatory variables	Levels (1)	Country-fixed effects (2)	Levels (3)	Country-fixed effects (4)	Levels (5)	Country-fixed effects (6)
Log income per capita (PPP in 1000 1985 $)	–0.104 –(0.69)	–0.160 (1.24)	–0.116 (0.77)	–0.194 (1.05)	–0.246 (1.12)	–0.0294 (0.15)
Income squared	0.0496 (0.80)	0.0772 (1.65)	0.0740 (0.90)	0.0955 (1.54)	0.105 (1.39)	0.0430 (0.69)
Year (–1900)	–0.0021 (0.79)	–0.0007 (0.27)	–0.0014 (0.33)	–0.0002 (0.05)	0.0012 (0.24)	–0.0029 (0.59)
Latin America	0.440 (5.62)	–	0.404 (4.50)	–	0.371 (3.97)	–
SW Asia	–0.0045 (0.07)	–	–0.0616 (0.57)	–	–1.01 (0.80)	–
ES Asia	0.0296 (0.46)	–	0.0404 (0.45)	–	0.0728 (0.76)	–
Africa	0.330 (2.48)	–	–	–	–	–
Income unit is person (or household)	–0.0602 (1.88)	0.0673 (2.03)	–0.0002 (0.01)	0.0855 (2.23)	–	–
Total expenditures (or income)	–0.0733 (1.77)	–0.0369 (1.14)	–	–	–	–
Disposable income (or pre-tax income)	–0.113 (3.72)	–0.0523 (2.16)	–	–	–	–
Constant	0.658 (4.23)	0.000 (0.0)	0.616 (2.55)	0.000 (0.0)	0.450 (1.57)	0.000 (0.0)
R^2	0.563	0.099[a]	0.484	0.071[a]	0.442	0.029[a]
Sample size	509	509	309	309	226	226
Mean dependent variable (standard deviation)	0.480 (0.265)	0.000[b] (0.120)	0.567 (0.273)	0.000[b] (0.137)	0.525 (0.236)	0.000[b] (0.110)
Joint significance on income coefficients $F(2, n)$ (p-value)	0.48 (0.49)	1.53 (0.22)	0.59 (0.44)	1.09 (0.30)	1.26 (0.26)	0.02 (0.88)

[a] Beneath regression coefficient in parentheses is the absolute value of the t statistic based on Huber (1967) standard errors that allow for heteroscedasticity of errors across countries.
[b] The country effects are not included in the R^2, because all variables are expressed as deviations from the mean of each variable for each country in the sample.

ity and the explanatory power of parallel regressions are similar. The estimates for the variance of the logs of income are reported in Table 3, which are subsequently used in the decomposition analysis (Eq. (1)). The regressions in Columns 1 and 2 include the maximum sized sample, first pooling all observations on levels, and then reestimating within countries, or equivalently including a fixed effect for each country. On the one hand, the fixed-effect estimates in regression (2) are not biased by the omission of time-invariant country-specific characteristics that affect income inequality and may be correlated with included control variables. On the other hand, the country fixed-effect estimates do not exploit the intercountry variation, which constitutes four-fifths of the variation in the pooled sample in regression (1). Regressions (3) and (4) in Table 3 are based on a restricted sample of 309 observations that includes predominantly gross household income data, but retains data on income distributed across persons for four countries for which there are no household data and which would otherwise be dropped from the sample: Argentina, Austria, China, and Yugoslavia. Regressions (5) and (6) rely on a sample of 226 observations based on only the preferred concept of income and recipient unit: gross household incomes. Control variables are added to capture (a) differences in the definition of the dependent variables, (b) the calendar time and stage of development (i.e., per capita income level), and (c) regional patterns.

The permanent income hypothesis, or most intertemporal models of consumption where utility is a concave function of consumption, suggest that inequality in total expenditures should be less than the inequality of income, because savings and transfers are expected to smooth consumption over time to increase the intertemporal discounted utility of income. On the basis of regression (1) the log variance in expenditures is accordingly about 15% smaller than the log variance in income (–0.0733/0.481), and within countries the log variance of expenditures is 8% smaller than that in income (–0.0369/0.481).

If the proportionate burden of taxes minus transfers is greater on the relatively rich than on the relatively poor, then such a progressive redistribution of income by the state would lead to a reduction in the log variance in net disposable income compared to that in gross income. The log variance is indeed reduced by about 23% (–0.113/0.481) when income is measured after taxes and transfers rather than pretax, but this gain is only half as large when estimated within countries (–0.0523/0.481).

It is more ambiguous how income inequality might differ if measured across households or across persons (with income), but, in these data, the personal income log variances tend to be substantially smaller in the pooled sample than the household log variances of income (–0.0602/0.481). Within-country comparisons suggest the opposite, however, that the log variance in incomes across persons is larger than that across households (+0.0673/0.481). Without a theory or a reliable procedure for relating the processes generating household and personal income distributions (cf. Deininger and Squire 1996), I am reluctant to mix data on households and persons, because it could conceal important regularities. The composition of households responds to income opportunities and therefore should be viewed as endogenous and possibly affected by urbanization, economic development, and possibly cultures. The third sample therefore relies only on data relating to gross household income (regressions (5) and (6)) to determine if parameters are sensitive to the exclusion of all data on the distribution of incomes across persons.

The most widely discussed empirical regularity in the distribution of personal incomes is the hypothesis advanced by Kuznets (1955, 1963) that modern economic growth in the now industrially advanced countries was associated with a reduction in the dispersion in personal incomes at the end of the 19th Century or in the first half of the 20th Century. Kuznets also speculated that there was an opposite tendency for the dispersion in personal incomes to increase at the onset of modern economic growth in the low-income countries, as labor is withdrawn from rural/agricultural activities and redeployed to more-unequal urban/nonagricultural sectors. This Kuznets inverted U-shape pattern in log variance (or Gini) in income with respect to economic development is not evident in these data collected from 1947 to 1995. Anand and Kanbur (1993) among others conclude that the traditional Kuznets pattern is weakened, eliminated, or reversed, when more recent and better data are analyzed with flexible functional forms. The linear term in income (PPP) per capita is consistently negative and the quadratic term is positive in these regressions accounting for the log variance (and for the Gini and Theil index). The last row in Table 3 reports that the quadratic parameters on the income variables are never jointly statistically significant. The lack of covariation between national income level and national household income inequality does not challenge the working assumption of the additive analysis of variance model in Eq. (1). Evaluated at the sample mean, a 10% increase in income per capita is associated with a 1.3% decline in log variance according to regression (2). The pattern of decreasing inequality with development appears to prevail within countries but reverses at higher income levels. [15] There is also some evidence of a downward trend in inequality over time, but this tendency is never statistically significant, implying an annual decline of 0.4% in the log variance, whereas within countries this trend is only one-third as large (regression (2)), unless the sample is restricted (regression (6)) to only data on household gross incomes.

Obviously, the effects of region cannot be estimated when individual country-fixed effects are included, since countries do not change their regional classification over time. In regression (1) on levels the log variances of incomes in Latin America are 0.440 higher than in the excluded high income group, or 91% above the sample means (0.440/0.481). In Africa the log variances are 0.330 larger than in the high income group. But in the case of Africa, the sample is small (six countries) and probably unrepresentative, whereas the deviant pattern of high inequality is well documented in Latin America (Deininger and Squire 1996). The two regions of Asia differ insignificantly from the high income countries.

Since four-fifths of the variance in measured income inequality in the pooled sample is “explained” by the country dummies, these estimated dummies are used to predict the log variance of household gross (before tax) incomes for the 56 countries in my maximum sample, allowing for the country’s income per capita and year effects to vary from 1960 to 1989 (according to regression (2), Table 3). As mentioned, these countries constitute 79% of the population in my sample of 120 countries as of 1960. For the remaining 64 countries, regression (1) is used to impute a value for the log variance of household gross incomes, based on the country’s income per capita, year, and region. The 3600 values of the predicted log variances of incomes by country and year are available from the author; they are summarized by region and selected years in Column 3 of Table 2.

It should be obvious that these estimates of intracountry household inequality are subject to a wide margin of error or uncertainty. For example, estimates for such large countries as China and India are not known with much precision, and estimates for 64, generally small, countries are imputed on the basis of only their national income per capita and region, because I lack two or more observations on their income distribution. Many of these gaps could be closed or an appraisal of the quality of the current data could be used to weight the regressions in Table 3. Moreover, Eastern Europe, China, and India may be experiencing in the last decade an increase in inequality associated with the reduced role of the state in the economy.[16] Better data might indicate that the intracountry average log variance in the world is thus no longer declining as appears to be the case in my sample.

6. Intrahousehold resource inequality

Barro and Lee (1994) have estimated adult education for 1960 to 1985 at 5-year intervals for 86 of the 120 countries in my working sample, but their data represent only 62% of its population, largely because they omit China, USSR, and Nigeria. Estimates by Dubey and King (1994) also construct gender-specific stocks of educational attainment for adults of different ages, using in addition lagged enrollment rates adjusted for mortality, but they include a smaller set of countries. Many assumptions are required by Barro-Lee to translate UNESCO completed/incompleted school attainment cross tabulations for adults into average years of schooling completed. Other assumptions are required to translate enrollment rates into attainments, such as completion and repetition rates. Moreover, both educational systems reporting school *enrollments* and respondents reporting their educational *attainment* to Censuses and Surveys may introduce distinctive errors that could be substantial and systematic by gender.[17] Until these potential inconsistencies in reporting gender differences in education are better understood, the estimates used here for specific countries should be treated with much caution, but perhaps regional and world trends will nonetheless be adequately summarized.

According to the Barro-Lee estimates of the years of schooling for adults age 25 or older, the population weighted world difference between the average education of men and women increased from 1.55 years in 1960 to 3.44 years in 1980, and only thereafter started to decrease. The log variance component attributable to these gender differences in adult schooling, given my working assumptions, increases from 0.0348 in 1960, to 0.0767 in 1980, and then decreases to 0.0566 in 1989, as summarized in Table 4. The convergence toward parity in the *ratio* of female to male educational enrollments noted in Schultz (1987, 1993) and Lichtenberg (1994) is apparently not sufficient to reduce the absolute gender gap in adult years of schooling until the 1980s.[18] Evidently, when most adult women in many African and some South Asian countries have received little or no schooling, the advancement of men's schooling in those countries first increases the gender gap in schooling for a time period before women start to catch

Table 4. Contribution to log variance of income due to gender differences in schooling[a] (weighted by total population)

	World total	High income countries[b]	Africa	Latin America	South Asia	East Asia
1960	0.0348	0.0116	0.0200	0.0120	0.0596	0.0676
1965	0.0417	0.0148	0.0276	0.0172	0.0696	0.0744
1970	0.0537	0.0168	0.0420	0.0208[c]	0.0976	0.0748
1975	0.0714	0.0192	0.0456	0.0176	0.1468	0.0744
1980	0.0767[c]	0.0212[c]	0.0588	0.0104	0.1536[c]	0.0768[c]
1985	0.0566	0.0192	0.0836[c]	0.0100	0.1036	0.0448

[a] Defined as $(r/2\rho)^2\ (E_f-E_m)^2$, where the rate of return to education, r, is assumed to be 0.15 and E_f and E_m are the average years of schooling completed by females and males age 25 or more, and the correlation of spouses schooling, ρ, is assumed to be 0.5.
[b] See row headings in Table 2 for definitions of regions.
[c] Maximum value over interval 1960 to 1985 within region.

up to that of men and close the absolute gender gap in average years of schooling. If these estimates are reliable, the turning point for women's education in the world occurred only in the late 1970s, and thereafter the reduction in the gender gap in schooling began to erode this important source of intrahousehold economic inequality.

Differences by region, shown in Table 4, are plausible. The gender gap in schooling is smallest in Latin America and largest in South Asia, and it may still be growing in Africa as of 1985. As noted earlier, if the returns on women's schooling, because it occurs disproportionately at the primary level, exceeds the returns on men's schooling, which is more concentrated at higher educational levels, this turning point in the magnitude of gender earning inequalities associated with schooling could have occurred earlier than estimated here. Regardless, the gender schooling component of the log variance of personal incomes appears to be a relatively small share of the sum of the intercountry and intracountry total inequality, as reported in Column 4 of Table 2. In 1960 it represents only 2.4% of the total, and by 1980 it had increased to 5.2%. Eliminating the gender gap in schooling entirely in the world, which could not occur for many years, would make only a modest direct contribution to reducing world inequality in personal incomes, according to these estimates. This finding buttresses the conclusion reached by a different route by Haddad and Kanbur (1990) concerning the relative unimportance of intrahousehold inequality in identifying the poor. Obviously, women reside in households in all income strata, more or less in the same proportion, except where female-headed households are especially common and concentrated among the poor. If, however, the gender inequality is more pronounced among the lowest income per capita countries, then my log variance decomposition or orthogonality assumption in Eq. (1) understates this source of personal income inequality. More research is needed to quantify the actual contribution of gender differences in education and health to personal differences in consumption, and their impact on welfare in particular countries and regions where gender inequality is most salient, as in South and West Asia and parts of Africa.

7. Conclusions and extensions for further research

Most analyses of the distribution of income focus on inequality across households within countries, the second of my three components in Eq. (1). Section 3 examines estimates of this second component of intracountry inequality in the last thirty years. Across countries the differences in income inequality are not strongly related to average income level or year, but there are salient regional differences in inequality, with income inequality being particularly large in Latin America and perhaps in Africa. There are also significant differences between countries, even within regions, that are not explained by income, year, or type of data. Four-fifths of the variance in the log variances (or Ginis) of income across countries and years are accounted for by country-fixed effects. Variation within countries over time are poorly accounted for by income changes and calendar time, e.g., R^2 is only 0.03 in regression (6) in Table 3. A first approximation of a country's inequality is then the average of all the observations for that country. There is nonetheless a weak tendency for the log variance of incomes to decline with time, and for there to be a U-shaped pattern of decreasing inequality as income increases for most poor and middle-income countries. The majority of the world's population, therefore, can expect to experience a small decrease in the intracountry variance in their log per capita incomes as they develop economically, but these trends are overshadowed by unexplained variation that may be real or an indication of the relative size of error in the measurement of inequality.

The first component of the log variance in incomes from Eq. (1) is obtained from real GDP estimates, based on either the conversion of local currencies according to their foreign exchange (FX) equivalence or according to their purchasing power parity (PPP). The differences in PPP income should accord with the consumption possibilities of persons more accurately than the FX incomes across countries at different levels of development. As expected, the log variance (or Gini) of intercountry income inequality is substantially smaller for PPP than for FX incomes. Also FX income levels and inequality are more volatile than PPP income levels and inequality, because foreign exchange crises and capital movements produce wider fluctuations in FX rates than PPP rates. PPP log variances in intercountry incomes increased by 25% from 1960 to 1968 and decreased by roughly the same amount from 1976 to 1985. The Gini concentration ratio based on intercountry PPP incomes increased about 6% from 1960 to 1968 and thereafter decreased about 6% by 1985.

Combining the intracountry and intercountry components of the log variance in household income inequality, Table 2 reports that total log variance in household incomes has changed relatively little over the entire 30-year period, rising in the first decade and declining modestly in the subsequent two decades, *if* we rely on the PPP conversion of local currency real income. Roughly two-thirds of this total approximation of the log variance of 1.5 is due to the intercountry component, and one-third is from the intracountry component.[19] But these shares differ markedly across regions. In Africa, Latin America, South Asia, and East Asia the intracountry log variances are two-thirds to three-fourths of the overall log variance in household incomes in the region. In contrast, in the high income country group the intracountry log variance component is initially in 1960 about half of the region's total log variance of household income, and thereafter the intracountry component increases to two-thirds of the total. This trend among the advanced countries is the obverse of the convergence in per capita income levels that is widely analyzed in the new economic growth literature (Dowrick and Nguyen 1989; Barro and Sala-i-Martin 1995).

The increasing economic productivity of women relative to men was thought to be a factor reducing the inequality of resources controlled by adult men and women within households. Based on empirical estimates of wage functions in which private wage returns to schooling for men and women are of a similar magnitude, a simple approximation of the gender gap in economic resources is derived as the square of the difference in years of schooling received by men and women weighted by the private return to schooling squared. This gender difference in schooling has, however, only begun to decrease after 1980, according to Barro-Lee estimates for a sample of 86 countries, and its expected contribution to the log variance in personal incomes is surprisingly modest. The closure in the gap between the schooling of men and women does not appear to be as powerful a lever to equalize the distribution of income within the household as expected. Another form of human capital that warrants more study as a source for the gender gap in productivity is health human capital. Life expectancy has increased more rapidly for women than men in this century, and has benefited differentially women relative to men within countries experiencing more rapid economic development (Schultz 1993). Finally, there is the increased participation of women outside of the home and their growing investment in skills that are rewarded primarily in the labor market. Although the gender gains in longevity can be viewed as with education, as a long-term social investment, the decision to participate in the labor force, within and outside of the family, is an individual choice that responds in the shorter run among adults to income opportunities, household incentives, tax and transfer policies, and social norms. A broader understanding of these reallocations in women's time from home production to the labor market will require explicit modeling of the determinants of fertility and the demographic transition itself.

Quality of data affects the confidence in conclusions

Without further summarizing the findings in this paper, it is appropriate to assess the confidence we can attach to the different levels in the analysis. The intercountry differences in income per capita are the product of many decades of work to construct and maintain on a comparable basis national income accounts around the world, and the margins of error should have been reduced over time as these methods have become more standardized and widely applied. The choice between the foreign exchange (FX) rates or purchasing power parity (PPP) rates to convert local currencies into dollar equivalents is shown to be important for measuring the levels and time trends in inequality. Conceptually, the PPP methodology is preferred, and the expected volatility in FX income is evident in the data. In general, therefore, the first stage of the analysis, based on PPP incomes, should capture reasonably accurately intercountry income inequality in the world.

The second level of the analysis consolidates estimates of intracountry inequality in gross incomes across households. At this second level the data are much less satisfactory, except for the majority of the high income countries and a few well-surveyed countries in Latin America and Asia. For most countries in Africa the imputed values and their change over time are only guesstimates. On the other hand, the African observations are not a large part of the population weighted sample, i.e., 11% in 1989. Intracountry inequality exhibits considerable heterogeneity across countries and strong persistence within countries. Changes within countries appear haphazard, as if due to measurement error, and are not correlated with growth in per capita income or time, according to the country fixed-effect estimates in Table 3. As a consequence, changes over time in intercountry inequality are likely to be of a greater magnitude than changes in intracountry inequality. This is illustrated by the reduced intercountry log variance in PPP income within the high income group from 1960 to 1980 of 0.49 to 0.24, which greatly exceeds the increase in intracountry log variance in PPP incomes from 1980 to 1989, estimated as from 0.43 to 0.44 (Table 2, Column 3). Better more recent data for the rapidly growing low-income and transition countries could modify conclusions regarding trends in intracountry inequality substantially (Gottschalk and Smeeding 1997a, b; Gustafsson and Johansson 1996).

The third level of the analysis seeks to quantify the contribution of gender differences in education as a factor determining intrahousehold inequality. The data are of recent vintage, with few consumers to evaluate their strengths and limitations, and therefore the empirical facts are subject to considerable uncertainty. The regional patterns in the gender-specific levels and changes over time in education appear plausible, and they do not imply major changes in aggregate inequality from this source. Much more research is needed on how to incorporate assortative mating and introduce health human capital before any firm conclusions are drawn on the magnitude of gender differences in personal welfare within households or its change over time.

Extending the time series on world inequality

National income estimates are not yet available from the Penn World Tables for most countries after 1989. However, the World Bank's *World Development Report for 1996* published 1994 GNP figures according to FX and PPP equivalent units for 106 countries out of my sample of 120.[20] The countries for which 1994 figures are available account for about 95% of the population in the original sample. In this restricted sample the level of PPP income per capita is 2.2% lower than for the full sample in 1989, and FX income is 0.2% lower. The log variance of PPP income per capita in 1989 for the restricted sample of 106 countries is 0.992, or about 1.9% lower than it was for the full 120 countries, whereas the log variance of FX income is 1.9% higher, respectively. From 1989 to 1994 the log variance of the PPP income per capita declined to 0.969 in the restricted sample, or by 2.3%, while the log variance of the FX income increased to 2.51, or by 1.9%. Thus, in the last 5 years the World Bank's GNP figures suggest that recent trends continued toward decreasing intercountry inequality as measured by the PPP income per capita, and increasing inequality according to FX income. In terms of the identical sample of 106 countries, this log variance of PPP income measure of inequality had essentially returned by 1994 to its 1960 value. In terms of the PPP Gini concentration ratio there is a negligible decline, from 0.550 in 1989 to 0.549 in 1994. Whether these 1994 GNP estimates from the World Bank are comparably constructed to those in the Penn World Tables remains unclear.

China's relatively rapid economic growth from 1989 to 1994 can account for the observed decline in world PPP inequality, as it had in the previous 15 years. Relatively, rapid population growth and below average initial income level in India added to world inequality, as measured by the log variance or Gini, based on PPP income per capita.[21]

The period before 1960 is more difficult evaluate, because the coverage and comparability of the sample of countries diminish. According to the estimates of Berry et al. (1983, Fig. 1), intercountry inequality in the 1950s declined slightly, as represented by either the Gini or mean logarithmic deviation of PPP income per capita. For the period 1960 to 1977, Berry et al. (1983) also show an irregular increase in intercountry inequality from 1961 to 1968, followed by the start of the decline in inequality that is documented more fully in this paper. Extending this approach back to 1870 might also be informative (Maddison 1989). Williamson (1996) concludes that the period of 1870–1914 was a period of international convergence, as was the period after the Second World War. But his analysis includes only higher income countries. It remains to be seen whether convergence is also occurring throughout the entire world in this early period. Basing comparisons over time on the winners introduces a serious bias. Although Pritchett (1996) may have failed to appreciate the current trend toward decreased inequality in PPP incomes since 1968, his conclusion that world inequality increased from 1870 to 1950 may well be sustained. This would make the trend in world inequality since 1968 all the more noteworthy. Of course, if the criteria for comparing economic inequality is through control over internationally-traded goods, the foreign exchange equivalent incomes should be used to define world inequality. Then such FX inequality has clearly increased, whether one refers to the Gini, log variance (Table 1) or Theil index (Table A-3).

Growth and distribution revisited

Why have economists focused mainly on inequality within nations? First, data on household and personal income are most readily collected within a national market area where variation in prices should be moderate and comparisons of income are a more satisfactory basis for inferring welfare. Second, some theories of economic growth suggest a tradeoff between income distribution and growth due to greater savings rates among the rich than poor (Kaldor 1956). More recently it has been postulated that imperfect credit markets could restrain the poor from making efficient levels of investment, particularly in non-collateralized human capital (Perotti 1993). Cross-country study of growth finds that countries that initially have less inequality in the distribution of income (or land) grow faster, holding constant for initial income per capita (i.e., convergence) (Barro and Sala-i-Martin 1992, 1995), and level of primary education (human capital) (Alesina and Rodrick 1994; Persson and Tabellini 1994). But this empirical regularity between inequality and subsequent growth is most frequently noted after 1960 and thus subsumes the period when Latin America and Africa report larger than average inequality and slower growth. The relationship between initial inequality and subsequent growth remains, nonetheless, robustly significant even when regional controls are included in the growth regressions (Clarke 1995; Birdsall et al. 1995; Bourguignon 1996).

One explanation for this empirical regularity is a "political economy equilibrium" in which a majority of the voters are more likely to favor growth-stimulating policies that encourage human or physical capital accumulation, when these forms of capital are already more widely distributed (e.g., Benhabib and Siegel 1992; Alesina and Rodrick 1994). But as Bourguignon (1996) notes, the empirical record does not confirm that less inequality stimulates higher investment rates in physical capital. Nonetheless, investments in human capital are not yet treated conceptually and empirically in the same consistent accounting framework that is applied to physical capital (Jorgenson 1995). The connection between inequality and human capital investments remains suggestive, if unproven.

Lucas (1988) hypothesized that the accumulation of human capital contributes more to economic growth than human capital earns as a productive factor in a competitive market. One way to distinguish such an increasing-returns growth externality associated with schooling would be to document spillover returns from schooling on aggregate growth that are not privately captured by individual workers in the labor market. Human capital externalities may be more important at the level of basic-primary education than at the level of advanced or technical vocational education, although empirical results are not in agreement on this point (e.g., Barro and Sala-i-Martin 1995). It is not implausible that such growth spillovers could also differ for male and female education, given the different tasks men and women perform and how they change during development.

Micro economic theories of household production and behavior have for three decades offered predictions for how the schooling of men and women would affect differently their allocation of time and their economic and demographic choices coordinated through families, such as marriage, fertility, and child rearing (Becker 1965; Schultz 1981, 1993). Some of the expected consequences of educating women and men have been viewed as generating a social externality, if, for example, they improve child health, increase intergenerational investments in human capital, and reduce population growth. Microeconometric studies of individual and family behavior have observed that women's schooling is partially correlated with lower child mortality rates, lower fertility, and smaller surviving family sizes in populations at widely different stages of development. Men's education is more weakly related to these same outcomes, and sometimes positively partially associated with fertility. This can be explained by a positive income elasticity of demand for children and the female time-intensity of producing children (Schultz 1981). Aggregate studies across countries, or within countries over time, document similar regularities between the schooling of young men and women of childbearing ages and these demographic outcomes (Schultz 1994): women's education is negatively associated with child mortality, fertility, and on balance with population growth, whereas men's education is less significantly related to all three and positively associated with fertility and population growth. These robust partial correlations, at both the micro and macro levels, between women's schooling and child health, fertility, and population growth suggest that increasing the share of human capital invested in women would hasten the demographic transition, facilitate further investments in the human capital of children, and through the resulting changes in age composition, increase domestic physical savings and investment.[22]

The specification of conventional growth regressions (e.g., Alesina and Rodrick 1994) can be readily modified to include initial schooling by gender. It is also possible, subject to the limitations of multicollinearity, to compare the estimated growth "effects" of different levels of enrollment and attainment for men and women. In regression (1) in Table 5 only primary enrollment rates are included for 84 countries with the necessary data, and the coefficient on male schooling is positive but substantially smaller than the coefficient on female schooling. The initial inequality, measured by the log variance in intrahousehold incomes in 1960, as derived above in Sect. 3, is inversely related to subsequent growth, as found in previous studies for smaller samples. Initial income is also inversely associated with growth controlling for initial primary education, confirming the general pattern of conditional convergence in this time period and sample (see notes to Table 5). Regression (2) includes also gender-specific enrollment rates at the secondary and tertiary level. Male secondary enrollments are significantly associated with subsequent growth, and a positive though insignificant coefficient is also obtained for male tertiary enrollments. But female enrollments at the secondary and tertiary level are insignificant and negative in sign. When all three enrollment rates are aggregated, the male and female coefficients in regression (3) are essentially identical in magnitude, although statistically different from zero only for women, $p<0.05$. These estimated growth patterns with regard to enrollment rates suggest primary educational levels for women may be particularly important along with secondary enrollment rates for men in forecasting growth from 1960 to 1989. However, in the regression (4) I use the Barro-Lee educational attainment series to replicate the puzzling finding noted earlier by Barro and Sala-i-Martin (1995; p. 431). It is unclear why there should be so little concordance between these two gender-specific educational attainment series.

Because labor market returns to primary schooling are of similar magnitude for men and women, my evidence that aggregate growth is more responsive to female than male primary school enrollments might be viewed as consistent with the hypothesis that primary education of women provides a growth externality. The challenge is now to analyze cross-country data over time to account for the *endogenous* determinants of growth (i.e., investment in physical and human capital) and the resulting international convergence (or divergence) in intercountry inequality. In such a growth framework, it will then be necessary to account simultaneously for the evolution of intracountry inequality, as affected by the level and distribution of human capital investments. These publicly subsidized and privately demanded human capital investments are observed to covary with gender inequalities. Progress will be needed to combine these mechanisms in a tract-

Table 5. Average annual PPP per capita income growth, 1960–1989 in percent, with gender-specific initial proxies for education[a]

Explanatory variables	(1)	(2)	(3)	(4)	Sample means (standard deviation)
Log variance of household income 1960	–0.0391 (5.05)	–0.0202 (2.17)	–0.0268 (3.03)	–0.0276 (3.40)	0.671 (0.224)
Log per capital GDP ($ 1000, 1985)	–0.00842 (3.39)	–0.0126 (3.91)	–0.0125 (4.06)	–0.0050 (1.47)	6.278 (0.864)
Enrollment ratios or investment flows 1960[b]					
Male primary	0.0147 (1.05)	0.0088 (0.65)			0.848 (0.207)
Female primary	0.0274 (2.88)	0.0263 (2.87)			0.758 (0.294)
Male secondary		0.0392 (2.32)			0.287 (0.230)
Female secondary		–0.0014 (0.08)			0.222 (0.220)
Male tertiary		0.0833 (1.13)			0.057 (0.065)
Female tertiary		–0.181 (1.62)			0.026 (0.037)
Male expected total			0.0024 (1.53)		7.20 (2.79)
Female expected total			0.0025 (1.85)		6.14 (3.02)
Estimated average years of educational attainment age 25+ (or stocks) 1960[c]					
Male				0.0109 (4.45)	3.90 (2.37)
Female				–0.0075 (3.03)	3.06 (2.55)
Constant	0.119 (7.50)	0.128 (7.04)	0.137 (7.85)	0.104 (4.95)	n. a.
R^2	0.416	0.496	0.404	0.395	
Dependent variable	0.0735 (0.0187)				
Sample size = 84					

[a] Sample includes 84 countries in Sample B of Appendix Table A-1, minus two countries: Malta and Guyana. The absolute value of t ratios is reported in parentheses beneath coefficients.

[b] Enrollment ratios are from UNESCO as reported in Barro-Lee (1994) data file. The expected total enrollment ratio multiplies the primary enrollment ratio by the duration of the primary system in years in 1984, plus the secondary enrollment ratio weighted by its duration, plus the tertiary enrollment ratio multiplied by the 5-year duration used to construct this ratio. This synthetic expected enrollment measure approximates the years of schooling completed by an average member of a cohort who experienced the current enrollment rates over their lifetime, and did not repeat any grades.

[c] Educational attainment is estimated by Barro and Lee (1994) from UNESCO cross-tabulations.

able manner that accounts for the basic features of world demographic and economic development and income distribution.

Endnotes

[1] It also precludes my using Theil's (1967) entropy index of inequality, which is additively decomposable but weights countries by income shares rather than population shares (Bourguignon 1979).

[2] Korzeniewicz and Moran (1997) examine income inequality in 46 countries (68% of the world's population) from 1965 to 1992, based largely on data from the World Bank's *World Development Report 1994*. They conclude that both the Gini and Theil T entropy Index of inequality in FX GNP per capita increased overall and between countries from 1965 to 1992, and most of this increase occurred after 1980. The between country Gini increased from 0.682 to 0.738 from 1965 to 1992, and the between plus within country Gini increased from 0.749 to 0.796. The total between and within T index of inequality increased from 1.15 to 1.32 in the same period. Since the Theil T index of inequality decomposes by income, they report that 79% of this total index of inequality is accounted for by the between country share of inequality in 1965, and 86% by 1992. The within country T index inequality thus decreased from 0.243 to 0.190. They conclude that within country inequality is a small fraction of the total and is not likely to influence trends in overall inequality. This was also the conclusion reached by Theil (1967) 30 years earlier in his study of world inequality applying his T index. Theil found that the total per capita income (FX) inequality was 0.530 in 1949 across 54 countries (48% of world population in 1950) and had decreased to 0.526 by 1957. The between country income weighted share of the total T index inequality was 86 and 88% in the two years analyzed by Theil (1967; Table 4.5). Theil also considered 1976, but on the basis of projections. The above values of Theil T inequality correspond to one-half of the variance of the logarithms of income, if personal incomes in the world are distributed lognormally (Theil 1967; p. 97). Jumping ahead, my estimates of the log variance of FX GDP per capita also increase from 1.63 in 1965 to 2.42 in 1989, roughly as do the estimates of Korzeniewicz and Moran (1997).

[3] Later versions of these Penn World Tables (Mark 5.6), however, include anomalous population figures for some countries. For example, Nigeria lost about one-fifth of its population from 1970 to 1971, and the populations of other countries are also affected for no clear reason in the Mark 5.6 version distributed by the NBER.

[4] I have considered here Gross Domestic Product (GDP), whereas it could be argued that Gross National Product (GNP) would be an appropriate measure of average welfare of national populations, which includes national income claims on foreign assets and excludes those domestic income flows owned by foreigners. Unfortunately GNP series are not available for the PPP series before 1970 and are available for only 85% of my sample population after 1973 when China is first included. In these 103 countries the population weighted GNP increases slightly more rapidly than GDP from 1973 to 1989 and the log variance of GNP decreases more rapidly than that of GDP. However, turning points and trends in intercountry GNP PPP inequality after 1973 parallel those reported here.

[5] Using the four points in the Lorenz Curve to estimate the Gini concentration ratio and the log variance will not yield a particularly precise estimate of these two parameters. In terms of the Gini, adding additional quantiles or points on the Lorenz Curve should increase inequality, and decreasing the size of the intervals from which the log variance is estimated is likely to increase the estimate of the log variance, for which Sheppard's correction is designed (Aitchison and Brown 1963). To preserve comparability I analyze my estimate of the Gini based on the same data I use to estimate the other measures of inequality, rather than rely on the Gini reported in the database. The Gini ratio estimated directly from the quantiles is correlated 0.97 with the Gini ratio reported in the database.

[6] Theil and Seale (1994) have also used the second index of entropy to measure intercountry PPP income inequality and decompose it into regional groupings of countries. Their Table 4 can be compared with my Table A-3. They also extend comparisons for some regions,

such as Europe, back to 1950, and thus quantify the intercountry convergence in incomes in Europe in this earlier decade.

[7] Decompositions of the Gini concentration ratio by underlying groupings of the population or factors explaining income have to be interpreted as though they were conditional on the overall value of the Gini. These decompositions are thus somewhat difficult to interpret. However, the log variance also violates the transfer axiom that is widely regarded as an attractive criteria on which to select a "natural" decomposition (Shorrocks 1980, 1984; Morduch and Sicular 1996).

[8] National income per capita from the Summers-Heston (1991) database refers to the arithmetic mean income that is logged in this analysis of intercountry income inequality. But the resolution of the variance in income in (1) refers to the logarithm of personal income, and thus the national income variable should refer to the mean of the logarithms of income.

[9] The distribution of households by their log incomes within a country tends to conform approximately to the lognormal. But the distribution of national log income per capita weighted by national populations are not normally distributed. Consequently, the log income at five quantiles of the world distribution of (PPP and FX) income per capita is reported in Table A-2 as they differ from the mean log income.

[10] Because the private rate of return on schooling tends to be higher in lower income countries in the world, the reported contribution of the gender differences in schooling to intrahousehold income inequality is probably understated in the less developed regions where returns are above the world average. Although there are fewer estimates of the assortative matching of spouses on schooling, i.e., ρ, from which to generalize across countries than there are estimates of the private wage returns to schooling, the pattern suggests ρ may increase with development and the level of education (Mare 1991; Shavit and Blossfeld 1993). Consequently, the assumption of a constant ρ may also relatively understate the contribution of gender differences in schooling to intrahousehold inequality in the least developed countries. A better empirical basis for imputing these intrahousehold variance components will likely increase the importance of gender inequalities in the overall distribution of personal income, particularly among the least developed countries of Africa and South and West Asia.

[11] Author's calculation based on the Personal Income and Expenditure Survey of Taiwan collected for 1995 for husband's age 30–34 and their wives. The correlation was about 0.6 for this age group in the 1976 round of the same survey. As implied by my formula, the variance in differences in years of schooling for husband's and wives is about three times larger than that one would expect based on the assumption of perfect assortative matching of spouses on schooling.

[12] Because the share of income invested in some countries is heavily determined by the state, and the state may allocate these investment funds among (state) enterprises in such a way as to yield relatively low rates of return, as in China and the USSR during the 1970s, it may be informative to base personal welfare comparisons on only consumption. But the national income accounts as reported by Summers and Heston (1991) do not distinguish between public sector allocations for investment purposes and for consumption purposes. Therefore, the only option is to analyze the narrower category of private consumption and thus ignore differences across countries in the share of GDP allocated to publicly provided consumption goods and services (Berry et al. 1983). The population weighted mean of the log of per capita private consumption in the world, converted to 1985 dollars at the PPP rate of exchange is about 0.40 lower than income in 1960, and 0.49 lower in 1989. The ratio of investment to GDP among the middle and lower income countries has increased compared to the ratio among OECD countries. The variance of the logs of PPP private consumption per capita is 0.851 in 1960, or 10% smaller than for income per capita. The variance in the log of per capita private consumption follows closely over time movements in the variance in log income, but after 1982 the PPP consumption log variance declines more slowly than it does for income, and in 1989 it is only about 2.6% lower than the income variance. Since growth in savings and investment ratios to GDP represents for many people an intertemporal reallocation of wealth that may contribute to the convergence in income and ultimately in consumption, there does not seem to be an overwhelming case for preferring the narrower private consumption series for assessing personal welfare levels

or inequality. Any trend toward less inequality after 1976 is, however, somewhat stronger in PPP income than in private consumption.

13 According to the PPP Gini measure of inequality, the effect of differential population growth is to increase intercountry inequality in income per capita after 1967 when they cross. Holding the initial year's population weights constant in 1960, the PPP Gini decreases slightly reaching in 1989 a level 2.5% lower than with current population weights, whereas the FX Gini increases 7% compared with the 14% increase for the series relying on the current population weights. If the population weights are fixed in 1989, the PPP Gini is again nearly constant, and it increases 11% based on the FX data.

14 The list of countries included in this sample is indicated in Appendix Table A-1 as sample Column A.

15 In regression (2) the turning point is 2819 per capita 1985 US$ and in regression (1) it is $ 2853.

16 For example, an analysis of wage, earnings and household income inequality in Czechoslovakia suggests substantial increases in inequality from 1983 to 1993 (Chase 1997) or in China (Morduch and Sicular 1996).

17 There have been relatively few validation studies to assess the numbers generated by both information systems, and develop consistency checks between them. National enrollment data by gender when lagged are often reasonably consistent with responses to "years of education completed" collected in representative surveys and censuses. But Barro-Lee adult schooling stock estimates are at times at odds in some countries with past enrollment patterns by gender.

18 The sample of 86 countries considered here (see Appendix Table A-1, Column B) is also larger than the 47 analyzed earlier by Schultz (1993).

19 Based on the Theil second entropy income inequality decomposition, the proportion of inequality associated with intercountry differences in per capita PPP income is also two-thirds. Compare Table A-3.

20 The countries lost from the sample for 1994 GNP are Angola, Gabon, Guinea, Iran, Malta, Myanmar, Puerto Rico, Reunion, Seychelles, Somalia, Syria, Taiwan, Yugoslavia, and Zaire. Algeria, Costa Rica and Hong Kong were reported in the *1995 World Development Report* with GNP by FX and PPP in 1993 US dollars. These were first inflated according to their real income growth from 1993 to 1994 in GNP per capita, and then converted from 1993 to 1994 dollars by the relevant US GDP deflator of 1.024. The World Bank does not provide sufficient documentation for these data to be confident that it is appropriate to chain together the Summer-Heston Penn World Table figures from Mark 5.5 for the 106 countries to those reported by the World Bank for 1994 or 1993. The changes reported here from 1989 to 1994 should, therefore, be regarded as tentative.

21 Removing China from the sample of 106 countries yields a larger log variance of PPP income of 1.124 in 1989 that does not change by 1994. Removing only India from the sample yields a log variance of PPP income that is 1.069 in 1989 and it would have decreased to 0.900 by 1994. The PPP Gini without China would have been 0.530 in 1989 and 0.554 in 1994, whereas without India in the restricted sample the PPP Gini would have been 0.535 in 1989 and 0.512 in 1994. According to either measure of inequality, China reduced world inequality and India increased it in this period.

22 However, the current state of evidence is not entirely convincing because it does not simultaneously model the marriage market and matching of spouses on unobserved characteristics (Foster 1996). If men with unobserved preferences for fewer children and greater intergenerational transfers (i.e., child quality) marry women with similar preferences that happen to be positively correlated with the schooling of these women, we could explain the widely noted correlation between women education and family outcomes, without relying on a market externality that justifies a differential subsidy favoring the education of women.

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Appendix

Table A-1. Countries included in world income sample from 1960 to 1989

	A.	B.		A.	B.		A.	B.
1 Algeria		x	47 Uganda		x	97 Korea, Republic	x	x
2 Angola			48 Zaire		x	100 Malaysia	x	x
3 Benin			49 Zambia		x	102 Myanmar		x
4 Botswana		x	50 Zimbabwe		x	105 Pakistan	x	x
5 Burkina Faso			52 Barbados		x	106 Philippines	x	x
6 Burundi			54 Canada	x	x	108 Saudi Arabia		
7 Cameroon			55 Costa Rica	x	x	109 Singapore	x	x
8 Cape Verde Islands			57 Dominican Republic	x	x	110 Sri Lanka	x	x
9 Central African Republic			58 El Salvador	x	x	111 Syria		x
10 Chad			60 Guatemala	x	x	112 Taiwan	x	x
11 Comoros			61 Haiti		x	113 Thailand	x	x
12 Congo			62 Honduras		x	116 Austria	x	x
14 Egypt			63 Jamaica	x	x	117 Belgium	x	x
16 Gabon	x		64 Mexico	x	x	119 Cyprus		x
17 Gambia			65 Nicaragua		x	120 Czechoslovakia	x	
18 Ghana	x	x	66 Panama	x	x	121 Denmark	x	x
19 Guinea			67 Puerto Rico			122 Finland	x	x
20 Guinea-Biss			71 Trinidad and Tobago	x	x	123 France	x	x
21 Ivory Coast	x		72 USA	x	x	125 Germany, West	x	x
22 Kenya		x	73 Argentina	x	x	126 Greece	x	x
23 Lesotho		x	74 Bolivia		x	128 Iceland		x
25 Madagascar			75 Brazil	x	x	129 Ireland	x	x
26 Malawi		x	76 Chile	x	x	130 Italy	x	x
27 Mali			77 Colombia	x	x	131 Luxembourg		
28 Mauritania			78 Ecuador		x	132 Malta		x
29 Mauritius	x	x	79 Guyana		x	133 Netherlands	x	x
30 Morocco	x		80 Paraguay		x	134 Norway	x	x
31 Mozambique		x	81 Peru	x	x	136 Portugal	x	x
32 Namibia			82 Suriname			137 Romania		
33 Niger		x	83 Uruguay		x	138 Spain	x	x
34 Nigeria			84 Venezuela	x	x	139 Sweden	x	x
35 Reunion			86 Bangladesh		x	140 Switzerland		x
36 Rwanda			88 China	x		141 Turkey	x	x
37 Senegal		x	89 Hong Kong	x	x	142 United Kingdom	x	x
38 Seychelles			90 India	x	x	143 USSR		
40 Somalia			91 Indonesia	x	x	144 Yugoslavia	x	x
41 South Africa		x	92 Iran		x	145 Australia	x	x
43 Swaziland		x	94 Israel		x	146 Fiji		x
45 Togo		x	95 Japan	x	x	147 New Zealand	x	x
46 Tunisia	x	x	96 Jordan	x	x	148 Papua New Guinea		x

A: Country included in intracountry inequality regressions in Columns 1 and 2, Table 3 (Deininger and Squire 1997).

B: Country included in Adult Education estimates by sex from Barro and Lee (1994).

Table A-2. Quantiles in the intercountry distribution of log income per capita, based on purchasing power parity and foreign exchange equivalence, 1960–1994

Year	Purchasing power parity income per capita						Foreign exchange income per capita					
	Differences between quantile log income and mean of log income					Variance of log incomes	Differences between quantile log income and mean of log income					Variance of log incomes
	10%	25%	50%	75%	90%		10%	25%	50%	75%	90%	
1960	–0.89	–0.89	–0.56	0.58	1.60	0.943	–1.06	–0.89	–0.81	1.22	1.79	1.511
1961	–1.10	–1.00	–0.53	0.65	1.67	1.050	–1.06	–1.06	–0.77	1.31	1.89	1.633
1962	–1.12	–1.00	–0.48	0.68	1.68	1.100	–1.15	–1.15	–0.73	1.27	1.96	1.723
1963	–1.09	–0.98	–0.48	0.65	1.66	1.075	–1.16	–1.16	–0.77	1.30	1.94	1.702
1964	–1.05	–1.03	–0.44	0.66	1.66	1.066	–1.11	1.11	–0.80	1.24	1.96	1.671
1965	–0.96	–0.96	–0.44	0.70	1.65	1.065	–1.02	–1.02	–0.82	1.22	1.97	1.634
1966	–0.87	–0.87	–0.44	0.74	1.65	1.088	–1.09	–0.94	–0.80	1.34	2.00	1.742
1967	–0.95	–0.95	–0.40	0.80	1.67	1.130	–1.13	–1.03	–0.81	1.49	2.01	1.819
1968	–1.05	–1.05	–0.38	0.86	1.71	1.199	–1.13	–1.11	–0.78	1.64	2.07	1.903
1969	–0.98	–0.98	–0.49	0.83	1.72	1.151	–1.15	–1.10	–0.75	1.65	2.10	1.868
1970	–0.93	–0.93	–0.46	0.84	1.71	1.107	–1.17	–1.06	–0.72	1.66	2.11	1.847
1971	–0.92	–0.92	–0.54	0.87	1.73	1.111	–1.17	–1.06	–0.75	1.66	2.16	1.890
1972	–0.92	–0.92	–0.66	0.90	1.76	1.158	–1.22	–1.06	–0.80	1.73	2.27	2.013
1973	–0.92	–0.92	–0.82	0.95	1.78	1.187	–1.24	–1.23	–0.82	1.77	2.34	2.055
1974	–0.94	–0.94	–0.65	0.98	1.76	1.180	–1.24	–1.17	–0.86	1.64	2.28	2.029
1975	–0.92	–0.92	–0.67	1.00	1.73	1.136	–1.34	–1.14	–0.85	1.61	2.44	2.023
1976	–0.98	–0.98	–0.69	1.05	1.76	1.192	–1.33	–1.24	–0.69	1.60	2.46	2.170
1977	–0.96	–0.96	–0.70	1.05	1.75	1.165	–1.30	–1.24	–0.61	1.59	2.43	2.151
1978	–0.93	–0.93	–0.65	1.05	1.73	1.143	–1.43	–1.43	–0.59	1.61	2.57	2.362
1979	–0.90	–0.88	–0.64	1.06	1.72	1.149	–1.38	–1.38	–0.71	1.53	2.54	2.341

Table A-2 (continued)

Year	Purchasing power parity income per capita						Foreign exchange income per capita					
	Differences between quantile log income and mean of log income					Variance of log incomes	Differences between quantile log income and mean of log income					Variance of log incomes
	10%	25%	50%	75%	90%		10%	25%	50%	75%	90%	
1980	–0.88	–0.81	–0.59	1.05	1.66	1.109	–1.37	–1.37	–0.52	1.44	2.51	2.256
1981	–0.86	–0.83	–0.54	1.07	1.66	1.071	–1.42	–1.42	–0.46	1.45	2.48	2.278
1982	–0.83	–0.83	–0.61	0.99	1.65	1.054	–1.36	–1.36	–0.38	1.23	2.45	2.208
1983	–0.80	–0.79	–0.64	0.90	1.62	1.029	–1.25	–1.25	–0.48	1.39	2.42	2.123
1984	–0.81	–0.72	–0.63	0.85	1.60	1.014	–1.24	–1.15	–0.49	1.39	2.42	2.102
1985	–0.81	–0.67	–0.63	0.80	1.60	0.970	–1.21	–1.07	–0.64	1.23	2.37	2.002
1986	–0.77	–0.67	–0.63	0.84	1.66	1.001	–1.19	–1.11	–0.87	1.32	2.68	2.224
1987	–0.74	–0.69	–0.64	0.78	1.69	1.019	–1.24	–1.24	–1.10	1.17	2.79	2.427
1988	–0.71	–0.71	–0.64	0.78	1.72	1.014	–1.19	–1.19	–1.06	1.14	2.86	2.435
1989	–0.70	–0.70	–0.60	0.71	1.74	1.011	–1.21	–1.13	–1.09	1.03	2.86	2.419
Restricted sample of 106 countries (see footnote 15)												
1989	–0.72	–0.66	–0.62	0.79	1.72	0.992	–1.21	–1.16	–1.11	1.08	2.85	2.465
1994	–1.01	–0.97	–0.97	0.43	1.72	0.970	–1.35	–1.35	–0.85	0.88	3.03	2.511

Table A-3. Theil reciprocal entropy inequality decomposition

Year	World sample			High income		Low income	
	Inter-country		Inter-household	Inter-country	Inter-household	Inter-country	Inter-household
	FX income (1)	PPP income (2)	(3)	PPP income (4)	(5)	PPP income (6)	(7)
1960	0.837	0.534	0.239	0.204	0.230	0.160	0.244
1961	0.886	0.574	0.239	0.190	0.228	0.194	0.244
1962	0.923	0.595	0.238	0.190	0.227	0.204	0.244
1963	0.907	0.594	0.237	0.188	0.225	0.193	0.243
1964	0.899	0.584	0.236	0.184	0.224	0.190	0.242
1965	0.889	0.583	0.235	0.183	0.223	0.185	0.241
1966	0.930	0.603	0.234	0.177	0.221	0.186	0.239
1967	0.969	0.618	0.232	0.168	0.220	0.199	0.238
1968	1.010	0.644	0.231	0.160	0.218	0.222	0.237
1969	0.995	0.626	0.230	0.157	0.216	0.207	0.236
1970	0.982	0.603	0.228	0.146	0.215	0.196	0.234
1971	1.002	0.606	0.227	0.139	0.213	0.200	0.233
1972	1.048	0.629	0.226	0.140	0.212	0.217	0.232
1973	1.056	0.641	0.225	0.136	0.210	0.233	0.231
1974	1.030	0.629	0.223	0.121	0.208	0.259	0.230
1975	1.032	0.607	0.222	0.111	0.207	0.247	0.228
1976	1.083	0.630	0.220	0.109	0.205	0.271	0.227
1977	1.076	0.619	0.220	0.108	0.204	0.261	0.226
1978	1.162	0.612	0.219	0.108	0.202	0.245	0.225
1979	1.155	0.614	0.218	0.109	0.201	0.256	0.224
1980	1.112	0.582	0.217	0.103	0.200	0.249	0.223
1981	1.127	0.578	0.216	0.104	0.200	0.247	0.222
1982	1.109	0.567	0.215	0.096	0.199	0.231	0.221
1983	1.050	0.559	0.214	0.097	0.199	0.210	0.220
1984	1.113	0.556	0.214	0.101	0.199	0.200	0.219
1985	1.086	0.535	0.213	0.101	0.200	0.186	0.218
1986	1.207	0.554	0.213	0.103	0.200	0.188	0.217
1987	1.304	0.562	0.212	0.103	0.201	0.196	0.216
1988	1.321	0.562	0.212	0.104	0.203	0.189	0.215
1989	1.311	0.563	0.212	0.105	0.206	0.188	0.213

Notes: The inequality in income across country per capita GDP based on foreign exchange rates increased according to the Theil (1967) inverse entropy index by 57% from 1960 to 1989, but the increase was only 5.4% when the more appropriate purchasing power parity price deflators are used. As with the log variance the Theil index reveals a sharp increase in PPP inequality until 1968, and a gradual decline from 1976 to 1985. The within country inter-household inequality declined about 12%, implying that the sum of the between country inequality components was 69% of the total in 1960, and was 73% in 1989, at the end of the world time series. Dividing the world into the high income (OECD plus Eastern Europe) and the low income (other), the intercountry inequality in the OECD declined by half from 1960 to 1982 (i.e., convergence), and increased only slightly thereafter, whereas the within country inequality declined by one-tenth, implying that the total inequality declined a quarter in the high income countries. In the low income group there was an increase in intercountry inequality until 1976 (i.e., divergence) and a gradual decline thereafter, whereas within country inequality may have declined slowly, leaving the sum of the two components of inequality approximately equal at 0.40 in 1960 and 1989. As noted in the text the movements over time and the shares of Theil inverse entropy inequality due to between-country and within-country inequality are very similar to those reported for the log variance in Tables 1 and 2.