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International Income Inequality: Measuring PPP Bias by Estimating Engel Curves for Food

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Abstract

Purchasing power adjusted incomes applied in cross-country comparisons are measured with bias. In this paper, we estimate the purchasing power parity (PPP) bias in Penn World Table incomes and provide corrected real incomes. The bias is substantial and systematic: the poorer is a country, the more its income tends to be overestimated. Consequently, international income inequality is substantially underestimated.

Our methodological contribution is to exploit the analogies between PPP bias and the bias in consumer price index (CPI) numbers. The PPP bias, and consequently real incomes, is measured by estimating Engel curves for food, which is an established method of measuring CPI bias. (*JEL*: D1, E31, F01)

1 Introduction

There are large differences between rich and poor people in the world. This is of major concern to economists, as well as to the public. Consequently, two major questions arise. First, how large are the differences between rich and poor, and second, do the differences get smaller or larger; i.e., do incomes diverge or converge across people and across countries? The answers depend on the measures used for comparisons. To illustrate, (per capita) income in China is six times larger if one uses Penn World Table (PWT) incomes rather than exchange rate based incomes (EX income).

In this paper, we study PWT incomes and estimate the bias in them by using Engel curves for food.[1] Furthermore, the relationship between the bias and the real income of a country is studied. Because the PWT produces purchasing power parity (PPP) adjusted incomes, the associated bias is referred to as the PPP bias. Having estimated the bias in PWT real incomes, we provide new estimates of real income. By comparing the estimated real incomes and the PWT incomes, the issue of how the bias influences estimated inequality and convergence is discussed. In addition, we discuss whether EX incomes provide better estimates of real income than do PWT incomes.

This paper incorporates five major findings. First, there are significant and substantial biases in the incomes published in the PWT. Second, there is a systematic relationship between the PPP bias and the real income of a country: the poorer is the country, the more its income tends to be overestimated. Third, the PPP bias causes a significant and robust underestimation of international inequality. The Gini index increases substantially when one corrects for the bias, and the distribution of estimated real incomes Lorenz dominates the distribution of PWT incomes. Fourth, for 22 countries between 1970 and 1995, predicted convergence is reduced when the PPP bias is corrected for. Fifth, EX incomes, which implicitly assume that PPP holds, provide better estimates of real incomes than do PWT incomes. Hence, in empirical work it is better to assume that PPP holds, than try to adjust for purchasing power differences by applying the PWT.

Although many macroeconomic studies rely on PWT data, few studies focus on the PPP bias in this data set. However, some contributors focus on one component of the bias, the so-called substitution bias, and use macro data to measure this bias (Dowrick

[1]The purchasing power parity bias is defined as the factor that converts real income into PWT measured income.

and Akmal, 2005; Hill, 2000; Neary, 2004; Nuxoll, 1994). In these studies, it is shown that, because of the substitution bias, international income differences tend to be underestimated by the PWT data. However, because there is another source of bias, known as the quality bias, the issue of underestimating international inequality cannot be robustly investigated without finding a way of measuring the overall PPP bias.

The main methodological contributions of this paper are twofold. First, our specific method based on Engel curve estimation enables estimation of the *overall* PPP bias, and enables the calculation of real incomes based on Engel curves (EC incomes). Second, applying micro data from household surveys eliminates the inaccuracies that arise from using aggregation techniques.

The difficulties of constructing PPP price indices are analogous to those of constructing consumer price indices (CPIs). A novelty of this paper is that it acknowledges and exploits this analogy by applying to the estimation of the PPP bias the method of Hamilton (2001) for estimating CPI bias. Engel curves for food are estimated by using micro data from different countries. Household incomes are made comparable by deflating household total expenditure by the macro price variable for consumption from the PWT. In the context of this method, we make the standard assumption that there is a stable relationship across countries between the budget share for food and real income; i.e., there is a unique Engel relationship for food in the world. Hence, any systematic difference in the estimated Engel relationship between a particular country and the base country (in our case, the United States) constitutes evidence of PPP bias for that country relative to the United States.

The paper is organized as follows. In Section 2, we discuss the causes of the bias and why the PWT tends to be systematically biased. In Section 3, we describe the empirical methodology in detail. In Section 4, we describe both the micro data and the macro price variables used. The analysis and main findings are presented in Section 5. In Section 6, for 22 countries, we discuss whether recent decades have seen the convergence or divergence of real incomes. In Section 7, we explain how EC incomes relate to EX incomes. Section

8 concludes the paper.

2 Explaining the Bias

PPP bias stems from two problems that are well documented in the price index literature, namely, the quality bias and the substitution bias (Costa, 2001; Hamilton, 2001; Hill, 2000; Neary, 2004). Most PPP calculations, including the Geary–Khamis calculations that underlie the PWT, are fixed-basket calculations. Fixed-basket calculations rely on using a set of homogenous goods, which generates the quality bias, and using a reference price vector for making comparisons, which generates the substitution bias. First, the quality of goods varies both over time and across countries. For example, it is not clear whether the observed price difference of cars between Poland and the United States reflects differences among the quality of the brands available in the two countries or represents a real price difference. Furthermore, some goods might be unavailable in some countries. For example, comparing the prices of Pakistani and Norwegian gur, which is a sugar substitute, is difficult simply because gur is not consumed in Norway. This is equivalent to the problem of quality differences because gur and sugar must be included in the same broad goods category, which makes it difficult to determine quality differences between these two goods correctly.

Second, the substitution bias arises because a reference price vector is applied to evaluate different countries' realized consumption bundles. Therefore, the fact that the consumers would have substituted their consumption away from relatively more expensive goods towards relatively less expensive goods, if faced with the constructed price level, is not taken into account.[2] Hence, unless consumers have Leontief preferences, both PPP and CPI measures incorporate substitution bias.

Both the quality bias and the substitution bias are systematic. First, because poorer countries tend to have products of lower quality than do richer countries, failing to adjust

[2]As shown in Appendix A, the Geary–Khamis price indices are Laspeyres indices as they compare each countries' price level with the constructed price level.

for quality causes poorer countries' incomes to be overestimated. Therefore, the difficulty of measuring quality can be expected to lead to poorer countries' incomes being systematically overestimated relative to richer countries' incomes.

Second, the greater is the difference between the own-price vector and the reference price vector, the greater is the substitution bias and, consequently, the greater the overestimation of a country's income tends to be (Nuxoll, 1994). Hence, using a vector of prices that are similar to those in rich countries, as the Geary–Khamis method does, causes incomes in poorer countries to be overestimated by more than incomes in richer countries; consequently, inequality is underestimated. (See Appendix A for further discussion and proof.)

3 Empirical Methodology

We use household micro data on nine countries to estimate national Engel curves for food, which represent the relationship between the household budget share for food and household real income. If two households, such as one in China and one in the United States, have the same PWT measured expenditure level and have the same demographic characteristics; i.e., the same age and numbers of children and adults, we attribute any difference in the budget shares for food to PPP bias.

There are several advantages of using food as the indicator good. First, because the income elasticity differs substantially from unity, the budget share is sensitive to bias in real income. Second, food is a nondurable good, which implies that expenditures in one period cannot provide a flow of consumption goods in another period. Third, studies for different countries and over different periods yield evidence that the Engel curve for food is log linear and stable, both over time and across societies (Banks et al., 1997; Beatty and Larsen, 2005; Blundell et al., 1998; Leser, 1963; Working, 1943; Yatchew, 2003). Fourth, food is arguably strongly separable from other goods in consumers' utility functions, which is necessary in order to decompose food and nonfood expenditures into a price index. (See Hamilton (2001) for an extensive discussion.)

3.1 Empirical framework – econometric specification

The standard almost ideal demand system (AIDS) specification (Deaton and Muellbauer, 1980) is:

$$m_{h,r,j} = a + b(\ln y_{h,r,j} - \ln P_j) + \gamma(\ln P_{f,r,j} - \ln P_{n,r,j}) + \theta X_{h,r,j} + \varepsilon_{h,r,j}, \tag{1}$$

where $m_{h,r,j}$ is the budget share for food, $y_{h,r,j}$ is nominal income, and $X_{h,r,j}$ is a vector of demographic control variables including the age of the household head and the numbers of children and adults in the household, for household h in region r in country j. P_j is the composite price of consumption in country j. $P_{f,r,j}$ is the price of food and $P_{n,r,j}$ is the price of nonfood items in region r in country j.

Because regional cross-country comparable price data are unavailable for the countries under study, the coefficient for relative prices, γ, cannot be estimated. Consequently, the main estimating equation excludes relative prices between food and nonfood items and, therefore, implicitly assumes that the budget share for food is unaffected by relative prices. However, because we have observations on national relative prices for five countries, a robustness check for relative price effects is conducted in the analysis described in Appendix B. When excluding the relative price effect, (1) can be simplified to:

$$m_{h,j} = a + b(\ln y_{h,j} - \ln P_j) + \theta X_{h,j} + \varepsilon_{h,j}. \tag{2}$$

Denoting the biased macro price variable for consumption given in the PWT P'_j and the PPP bias for country E_j, the unbiased price variable P_j, can be expressed as:

$$P_j = P'_j * E_j. \tag{3}$$

Equation (2) can therefore be expressed as:

$$m_{h,j} = a + b(\ln y_{h,j} - \ln P'_j) + \theta X_{h,j} + \sum_{j=1}^{N} d_j D_j + \varepsilon_{h,j}, \tag{4}$$

where D_j is the country dummy representing the PPP bias. More specifically, the country dummy coefficient, d_j, is a function of the PPP bias, E_j, and the coefficient for the logarithm of real income, b, as follows:

$$d_j = -b \ln E_j. \tag{5}$$

The specification represented by (4) is our preferred specification. Hence, the PPP bias is:[3]

$$E_j = e^{-\frac{d_j}{b}}. \tag{6}$$

Because the budget share for food is decreasing in real income (i.e., b is negative), the estimated bias exceeds unity if the estimated country dummy coefficient is positive.

According to (3), dividing the PWT incomes by the PPP bias gives new estimates of real incomes; i.e. the EC incomes. If the bias exceeds unity, the PWT consumption price is underestimated and, therefore, the real income of the country is overestimated. The larger is the estimated country dummy coefficient, the larger is the estimated bias, and consequently, the more is national per capita real income overestimated.

[3]For robustness analysis, alternative specifications were estimated: see Appendix B.

4 Data

Micro data are used to estimate the Engel curves to determine the bias in the PWT macro price variable. We discuss the micro data and the macro price variables in Sections 4.1 and 4.2, respectively. In Section 4.3, we discuss the UN data used in Section 5.3 and Sections 6 and 7.

4.1 Micro data from household surveys

The full sample comprises observations on 52,543 households from nine countries. Table 1 provides an overview of the different surveys. The household data for Azerbaijan, China, Nicaragua, and Côte D'Ivoire are from the World Bank's living standard measurement surveys (LSMS). The data for the United States are from the Consumer Expenditure Surveys and the United States Bureau of Labor. The Hungarian data are from the Hungarian Central Statistical Office (Household Budget Survey Section). Luxembourg Income Studies (LIS) provided the data for France, the United Kingdom, and Italy.[4] The nine countries selected publish nationally representative surveys and represent a geographical spread that includes high- and low-income countries.[5]

Because it is difficult to harmonize data from different surveys, our analysis relies primarily on sources that present harmonized analyses, such as the LIS and, to a lesser extent, the LSMS. The choice of estimation technique is limited by the lack of panel data on lower-income countries. In addition, data limitations for some countries restrict the choice of explanatory variables.

To estimate the preferred specification, we include only households with two children and two adults. Hence, we exploit an advantage of micro data, which is that they can be used to analyze households of the same composition and size to avoid the inaccu-

[4]Detailed information on different LSMS and LIS studies can be found on the World Bank and LIS websites, respectively (Luxembourg Income Studies, 2006; World Bank, 2005).

[5]All data are nationally representative except those from China. For China, no national representative study is available. The Chinese data include households from the provinces of Hebei and Liaoning, which implie that only rural households are covered.

racies generated by heterogeneous household composition. For robustness analysis, we estimated equations based on the whole sample.

[Table 1 about here.]

Many of the households included in the sample are farm households, for which home-produced food accounts for much of total household consumption. We account for this by incorporating expenditure on home-produced goods in the expenditure variables.

4.2 Macro price variables

In the standard AIDS specification, three macro price variables are included. The first, P'_j, is a composite price index for all consumption goods in country j, which is constructed by using the Geary–Khamis method, and presented in the PWT. The other two macro price variables are the composite price index for food items, $P_{f,r,j}$, and the one for nonfood items, $P_{n,r,j}$.

Because the household surveys are conducted in different years, the macro price variable for consumption in the PWT relates to different years. Because the consumption price in the PWT is reported in current prices, we use the United States exchange rate and CPI to make income levels comparable across countries and time. The macro price variable for consumption and the exchange rate are taken from Penn World Table 6.1 (Heston et al., 2002). The United States CPI is taken from the World Bank's World Development Indicators online (World Bank, 2007).

Because of a lack of data, the preferred specification (represented by equation (4)) does not include relative prices between food and nonfood items. Unfortunately, there are no cross-country regional price data for food and nonfood items. Few countries report regional price variations, and those that do report them do so relative to a base year. The price in one region is compared with the price level of that same region in a different year. Therefore, these data cannot be used for cross-regional comparisons between specific years. The same applies to national price indices; e.g., the food price index produced by the World Bank. Because these indices represent price levels relative to a base year, they cannot be used to compare relative prices across countries.

Cross-country comparable *national* prices for food and nonfood items for 1980 (phase IV) from the International Comparison Project are reported by Neary (2006). Combining these data with the price indices from the World Bank yields comparable *national* relative prices for Hungary, the United States, France, the United Kingdom, and Italy. However, because we have no regional price data, the coefficient of the relative price cannot be identified. To overcome this problem, in Section 5, we use Costa's (2001) estimated coefficient for relative prices, γ, for a robustness check. Using Costa's estimated coefficient enables inclusion of *national* relative price levels for these five countries. In this way, estimation incorporates the relative price effect.

4.3 UN data used in the extended model

We extend our analysis based on the estimated Engel curve for food and by using mean household data from the UN Statistics Division (Common Database). Nine hundred and eighty-three observations on mean household consumption and budget shares are included, covering 46 different countries from 1970 to 1995. We use data on final household expenditure in national currencies at current prices (Series code 21650). To make final household consumption comparable among countries and across years, we use the PWT price of consumption, the exchange rate, and the United States CPI. For the main model, these data are taken from Penn World Table 6.1 (Heston et al., 2002) and the World Bank's World Development Indicators online (World Bank, 2007).

To obtain the household mean of the logarithm of income from mean household income, we use the distributions estimated by Sala-i-Martin (2006).[6] Information on demographic controls are also obtained from the UN. The numbers of children and adults and thus the OECD's adult equivalence scaling, can be calculated directly (Series code

[6]In addition to the mean of household expenditure, information on the distribution is required because the mean of the logarithm of x is not equal to the logarithm of the mean of x ($ln(mean(x)) \neq mean(ln(x))$).

13681 and 1070). The age of the household head is approximated from the relationship between life expectancy and the age of the household head. This is estimated by using data from the nine household surveys that provide the background for the main analysis of this paper and from data on life expectancy from the UN statistics division (Series code 13630).

In Section 6, we illustrate the effect of the PPP bias by calculating convergence for 22 countries based on PWT incomes and the adjusted estimates of real income (EC incomes). The countries for which we have observations are Belgium, Canada, Denmark, Ecuador, Finland, France, Hong Kong, India, Iran, Ireland, Israel, Italy, Japan, the Republic of Korea, Mexico, Norway, Singapore, South Africa, Sweden, Switzerland, Thailand, and the United States. There may be a selection problem because data may be more likely to be available on richer countries and countries that have higher growth rates. Hence, this analysis is merely illustrative, and does not constitute an extensive analysis of the extent to which rich and poor countries converge or diverge.

The EX incomes are constructed by using the exchange rates from Penn World Table 6.1 (Heston et al., 2002), which are used to measure final household consumption from the UN (UN Statistics Division, Common Database, Series code 21650) in a common currency. The EX incomes are compared to the EC incomes in Section 7.

5 Analysis and Findings

In this section, two models are estimated and two data sets are analyzed. First, the PPP bias is estimated by using household surveys from nine countries, and the findings from this model are discussed in detail. Second, an extended model is estimated, and aggregated UN household data are used to measure the PPP bias for other countries and for other years. Then we discuss whether our findings can be generalized to the nine countries.

5.1 The main model based on household surveys

[Table 2 about here.]

The regression results are presented in Table 2. The preferred specification estimates equation (4) on the subsample of households with two children and two adults. In Appendix B, for a robustness check, this model is estimated on the sample of all households; the main results are unaffected. The estimated income elasticity for food is similar to those from related studies (Costa, 2001; Hamilton, 2001; Beatty and Larsen, 2005; de Carvalho Filho and Chamon, 2006). By construction, the United States country dummy coefficient is equal to zero, whereas all the other dummy coefficients are used to measure the PPP bias when comparing incomes with the United States. All these are significantly different from zero. All countries except for the United Kingdom have a positive dummy coefficient; i.e., the macro price variables in the PWT underestimate the macro price levels relative to the United States macro price level. Therefore, all countries' real incomes, except for those of the United Kingdom, are overestimated relative to United States income in the PWT. The estimates also show that the non-OECD countries, China, Nicaragua, Azerbaijan, and Côte D'Ivoire, have substantially higher dummy coefficients than the OECD countries.

The dummy coefficients indicate substantial PPP biases. Our first major finding is that the biases in PWT incomes are substantial. Côte D'Ivoire has the largest bias and its real income is overvalued by a factor of seven. China's dummy coefficient indicates that its real income is overvalued by a factor of five in the PWT.

[Figure 1 about here.]

Figure 1 reveals our second major finding: there is a negative relationship between the PPP bias and real income levels. That is, the estimated bias is much higher for poorer countries. The interpretation of this finding is related to the discussions of Section 2 and Appendix A. As expected, the larger is the overall bias the poorer is the country. Moreover, our estimates of the bias exceed those of other studies that focus on the substitution bias (such as that of Nuxoll, 1994). This indicates that not only the substitution bias, but also the quality bias systematically leads to an overestimation of poorer countries' incomes. The finding that Hungary has a substantial positive bias supports this conclusion. Hungary has prices that most resemble the prices used in the PWT comparisons (Nuxoll, 1994). Hence, for Hungary, one would expect the substitution bias to be small. Thus, our finding that Hungary has a substantial positive bias gives additional support to the predicted direction of the quality bias.

The third major finding is that international inequality, as measured by the Gini index, is substantially underestimated. Table 3 reports different estimates of international inequality based on the Gini index. The table shows that the index increases substantially when the correcting for the PPP bias for both the unweighted and population weighted cases. The unweighted Gini index increases from 0.45 to 0.58 after correcting for the bias, and the population weighted Gini index increases from 0.58 to 0.73.[7]

[Table 3 about here.]

[Figure 2 about here.]

It is important to examine whether the estimated increase in inequality is robust. That is, does inequality increase if other inequality measures are used, or is the use of the Gini index essential for our findings? Figure 2 presents the Lorenz curves for the PWT and the EC incomes, and forms the basis of our fourth main finding: the distribution of EC incomes Lorenz dominates that of the PWT incomes. Hence, we obtain the robust finding that inequality is underestimated in the PWT.[8]

5.2 An extended model

Figure 1 shows a negative relationship between the PPP bias and the EC incomes. However, the relationship is based on only nine countries, which comprise rich and poor ones.

[7]For a general discussion of these inequality concepts, see Milanovic (2005).

[8]All inequality measures that satisfy the Pigou–Dalton criterion, which is considered uncontroversial, support this conclusion (Fields and Fei, 1978; Sen, 1997).

The question is whether we can generalize this finding to other countries and to other years. Unfortunately, for most countries we do not have data that are sufficiently detailed to identify the bias and the EC incomes. However, we can extend the model by using the nine countries as benchmark countries and then add means of variables for other countries and for other years.

Deviations in a country's relationship between the mean household budget share and the mean household PWT income from the estimated Engel curve can be utilized to determine the bias in the PWT consumer price for that country. From equation (2), it follows that:

$$\overline{m_k^t} = a + b\overline{\ln y_k^t} - b \ln P_k^t + \theta \overline{X_k^t}, \tag{7}$$

where $\overline{m_k^t}$, $\overline{\ln y_k^t}$, and $\overline{X_k^t}$ are the mean household budget share for food, the mean household logarithm of nominal consumption, and the mean household demographic characteristics in country k at time t, respectively.

We measure the overall price of consumption in country k at time t by:

$$\widehat{\ln P_k^t} = \frac{\widehat{a} + \widehat{b}\widehat{\overline{\ln y_k^t}} + \widehat{\theta}\overline{X_k^t} - \overline{m_k^t}}{\widehat{b}}, \tag{8}$$

where $\widehat{a}$, $\widehat{b}$, and $\widehat{\theta}$ are the estimated coefficients in the benchmark model and $\widehat{\overline{\ln y_k^t}}$ is the estimated mean household logarithm of nominal expenditure.

The PPP bias for country k at time t is measured indirectly by using the estimated price ($\widehat{\ln P_k^t}$) and the PWT prices (P') (cf. Equation (3)):

$$E_k^t = \frac{\widehat{P_k^t}}{P'} \tag{9}$$

5.3 Generalizing the results

By applying the method described in the previous section, we estimate the PPP bias for 46 countries in different years; we use 983 observations in total. We pool all observations, and as shown in Figure 3, our first major result is preserved: the poorer is the country, the larger is the bias. As shown in Table 3, for this larger group of countries, estimated inequality increases substantially after correcting for the PPP bias.[9]

[Figure 3 about here.]

Because the distribution of EC incomes Lorenz dominates the distribution of PWT incomes, our finding of underestimation is robust. Thus, the extended analysis shows that we can generalize the findings from the benchmark countries.

6 Convergence or Divergence? An Illustration

There are a number of important implications of substantial and systematic PPP bias. One is whether countries converge or diverge. To illustrate this, we study the question of convergence for 22 countries by analyzing the change in inequality from 1970 to 1995.

Table 4 reports the Gini index for these two years. The table shows that convergence is much higher when based on PWT incomes rather than EC incomes; i.e., the decrease in the Gini index is smaller after correcting for the PPP bias in the PWT incomes.

[Table 4 about here.]

Figure 4 displays the Lorenz curves for the uncorrected PWT measure, whereas Figure 5 displays the Lorenz curves for the corrected PWT measures. Applying the uncorrected measure yields a robust result for convergence, because the 1995 distribution Lorenz dominates that of 1970. This strong conclusion does not hold after correcting for the PPP bias because the Lorenz curves cross in this case (see Figure 5).

[9] We removed the 10 percent of countries with the highest bias and the 10 percent with the lowest bias. Our analysis focuses on the 80 percent of observations in the middle of the distribution. The results are strengthened when the extreme observations are included.

[Figure 4 about here.]

[Figure 5 about here.]

7 How do the EC Incomes Relate to the EX Incomes?

Traditionally, international comparisons of income relied on the exchange rate based method, which involves simply transforming incomes into a common currency, such as the United States dollar. However, if PPP does not hold, and if prices for nontraded goods differ between countries, then using the exchange rate yields biased estimates of real incomes. Given that PWT incomes are biased, an interesting empirical issue is which approach provides the best estimates of real income. To approach this question, we first examine China. China is 27 percent poorer if EX, rather than EC, income is used to measure its real income. However, our analysis shows that China is more than 400 percent richer based on its PWT income rather than on its EC income. Hence, China's EC income is much closer to its EX income than to its PWT income.[10] Turning to other countries and to estimated international inequality, we find that the EX incomes provide better estimates of real incomes than do the PWT incomes. Figure 6 shows the Lorenz curves for the EX incomes, the PWT incomes, and the EC incomes, based on the extended model. Both the EX incomes and the EC incomes Lorenz dominate the PWT incomes. The Lorenz curve for the EX incomes is close to the Lorenz curve for the EC incomes. However, it is further away from the diagonal for the lower tail of the distribution. This indicates that applying the EX method underestimates the incomes of the poorest countries. This may be because prices are lower in these countries, even if they are fully adjusted for quality.

[Figure 6 about here.]

Hence, despite the empirical evidence against PPP, it seems better to assume that PPP does hold by using the EX incomes, than to apply PWT incomes.

[10]The considerable overestimation of China's income implied by the PPP measures of the World Bank has recently been acknowledged and adjusted for (The Economist (2007a and 2007b; Keidel (2007)). In this paper, we show that this problem also applies to the PWT-based PPPs.

8 Concluding Remarks

In this paper, we used household surveys from nine countries and UN mean household data to provide initial estimates of the overall purchasing power parity (PPP) bias in the Penn World Tables (PWT). Although the PWT incomes are extensively used by economists, there are few studies that investigate the bias in these measures. We found evidence of substantial and systematic bias, and provided a clear interpretation of the source of this bias. Because of substitution bias and quality bias, poorer countries' incomes are overestimated relative to those of richer countries. Consequently, the PPP bias causes a significant and robust underestimation of international inequality.

We illustrated how the PPP bias affects estimated convergence among 22 countries between 1970 and 1995. Estimated convergence is reduced after correcting for the PPP bias. The finding of convergence is robust for the PWT incomes. However, robustness does not hold after correcting for the PPP bias.

The PPP bias is so substantial that applying the traditional exchange rate based method, which implicitly assumes that PPP holds and that prices for nontraded goods do not differ among countries, yields better estimates of real incomes and international inequality.

Several robustness checks, reported in Appendix B, show that the main findings are not driven by the misspecification of functional form, differences in relative prices, or household composition. However, this study, as well as other studies based on micro data (or macro data based on micro data), could have benefited from a greater availability of already harmonized studies. It would be interesting to base future research on more detailed data than those currently available. There are two reasons for this. First, if panel data were available for poor countries, as is the case for OECD countries, more sophisticated estimation techniques could be used. Second, the availability of harmonized data for rich and poor countries would facilitate cross-country comparisons based on micro data, and consequently, more than nine countries could be used to estimate the Engel curve.

Section 2 and Appendix A provide explanations of the bias, and show that the bias is systematic, as our empirical analysis suggests. In addition, Dowrick and Akmal (2005) and Nuxoll (1994) provide theoretical support for our empirical finding that income differences are underestimated in the PWT. Future research could be used to generalize these implications for the Geary–Khamis method.

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Appendix A The Substitution Effect

> ... the direction and magnitude of bias in GK (Geary–Khamis) bilateral income ratios depends on whether the GK price vector corresponds most closely to the relative price structures of high income (high productivity) countries, in which case most bilateral ratios will be underestimated, or whether the GK price vector corresponds most closely to the relative price structures of low income (low productivity) countries in which case most bilateral ratios will be overestimated. The former situation is **most likely** to apply given that the GK method weights each country's price vector by its share in total GDP, implying that more weight is given, ceteris paribus, to the price vectors of the richer countries. [Dowrick and Akmal, 2005, our emphasis.]

In this appendix, we show that it is not only **most likely** that the Geary–Khamis prices correspond more closely to the relative price structure of richer countries, but that this follows by construction.

A.1 Geary–Khamis constructed world prices

The Geary–Khamis method, underlying the PWT, calculates a vector of reference prices applied for comparison $\Pi = [\Pi_1, \Pi_2, ..., \Pi_m]$ (referred to as the GK price vector above). The main question that we study is whether the reference price vector resembles the prices of richer countries more than those of poorer countries. That is, we investigate whether the price vector of a rich country j, $[p_{j1}, p_{j2}, ..., p_{jm}]$ is given a greater weight when constructing the reference price vector Π, than is that of a poorer country, f, $[p_{f1}, p_{f2}, ..., p_{fm}]$. The reference price for good i, is given by:

$$\Pi_i = \sum_{j=1}^{N} \left(\frac{q_{ij}}{\sum_{j=1}^{N} q_{ij}} \frac{p_{ij}}{P_j}\right), \tag{10}$$

where q_{ij} is the quantity of good i consumed in country j, p_{ij} is the price of good i in country j, P_j is the overall price index of country j, and N is the total number of countries in the system. This equation suggests that the Geary–Khamis method gives a greater weight to richer countries' prices. To see this, consider the following derivative:

$$\frac{\delta \Pi_i}{\delta \frac{p_{ij}}{P_j}} = \frac{q_{ij}}{\sum_{j=1}^{N} q_{ij}}. \tag{11}$$

This derivative increases in country j's share of the total consumption of good i; i.e., the derivative is larger the richer is this country in terms of the consumption of this good. However, the overall price level of country j, P_j, is endogenous and is given by:

$$P_j = \frac{\sum_{i=1}^{m} p_{ij} q_{ij}}{\sum_{i=1}^{m} \Pi_i q_{ij}}. \tag{12}$$

When considering this indirect effect, the problem turns out to be more complicated. To simplify, we study the special case of two countries and two goods. By showing that the cross derivative of the reference price with respect to quantity and price is positive; i.e., that the richer is a country, the greater is the weight given to its price, we show that richer countries are given a larger weight in the construction of the reference prices.

In the specific case of two countries and two goods, one needs to determine whether the two prices in a rich country, say country 1, are given a greater weight in the construction of the two reference prices than are those in a poorer country, country 2. This is investigated by comparing the influence of country 1's prices on the reference vector Π when countries 1 and 2 are equally rich with the corresponding influence when country 1 is richer. That is, when both countries are equally rich, in the sense that they consume equal amounts of both goods, what happens to the weight as country 1 gets richer? Country 1 can get richer in three different ways. First, it can start consuming more of both goods. Second, it can start consuming more of only one of the goods. Third, it can start consuming more of one good and less of the other good, but in such a way that the total expenditure level in the country is higher than that of the other country,

$(\Pi_1 + \Delta\Pi_1)\Delta q_{11} + (\Pi_2 + \Delta\Pi_2)\Delta q_{21} > 0$, with q_{ij} being the quantity of good i consumed in country j. The third case can be reformulated so that it can be analyzed in the same way as the second case. The results for the first case are shown below. The main results for the second case are the same as those for the first case and thus are not reported.

The weight that a specific price p_{ij} is given in the construction of a specific reference price, Π_i, is given by the derivative of the reference price with respect to that price:

$$w_{ij} = \frac{\delta \Pi_i}{\delta p_{ij}}. \tag{13}$$

If w_{ij} is larger the richer is country j, then the richer country gets a higher weight in the construction of the Geary–Khamis prices underlying the PWT.

Focusing on the weight in the construction of the reference price of good 1, we must find the change in the weight from the direct effect of p_{1j} and from the indirect effect of p_{2j} when country j gets richer. For this analysis, the countries are assumed to be equally rich initially, and then country 1 gets richer in the following way:

$$\Delta q_{11} = \Delta q_{21} = a > 0$$

.

The parameter a is a small positive constant. If country 1 has a higher weight in the construction of the reference prices after this change, then the richer country gets a higher weight in the construction of the Geary–Khamis reference prices. Hence, whether country 1's price is getting a greater weight in the construction of the reference price vector can be determined by finding the sign of the following derivative:

$$a\left(\frac{\delta^2 \pi_1}{\delta q_{11} \delta p_{11}} + \frac{\delta^2 \pi_1}{\delta q_{21} \delta p_{11}}\right). \tag{14}$$

In the two-country two-good case, the Geary–Khamis system, given by equations (9) and (11), reduces to:

$$\Pi_i = (\frac{q_{i1}p_{i1}}{P_1} + \frac{q_{i2}p_{i2}}{P_2})(q_{i1} + q_{i2}), \tag{15}$$

where q_{ij} is the quantity of good i consumed in country j, p_{ij} is the price of good i in country j, and P_j is the overall price index of country j, which is given by:

$$P_j = \frac{p_{1j}q_{1j} + p_{2j}q_{2j}}{\Pi_{1j}q_{1j} + \Pi_{2j}q_{2j}}. \tag{16}$$

In this framework, the weight of country 1's price of good 1 in the reference price of good 1 is given by:

$$w_{11} = \frac{\delta\pi_1}{\delta p_{11}} = \frac{q_{22}q_{21}p_{21}q_{11}(q_{11}p_{12}q_{12}q_{21} + q_{11}p_{12}q_{12}q_{22} + q_{21}p_{22}q_{22}q_{11} + q_{21}p_{22}q_{22}q_{12})}{(p_{11}q_{11}q_{12}q_{21} + p_{11}q_{11}q_{12}q_{22} + p_{21}q_{21}q_{11}q_{22} + p_{21}q_{21}q_{12}q_{22})^2}. \tag{17}$$

The weight of country 1's price of good 1 in the construction of the reference price of good 1 is then given by:

$$\begin{aligned}
&\Delta\frac{\delta\pi_1}{\delta p_{11}}|_{q_{11}=q_{12},q_{21}=q_{22}} = a(\frac{\delta^2\pi_1}{\delta q_{11}\delta p_{11}} + \frac{\delta^2\pi_1}{\delta q_{21}\delta p_{11}})|_{q_{11}=q_{12},q_{21}=q_{22}} \\
&= a(q_{11}^3 p_{12}q_{12}q_{22}p_{11} - q_{11}^3 p_{12}q_{22}p_{21}q_{21} + q_{11}^3 q_{21}p_{22}q_{22}q_{12} + q_{11}^3 p_{12}q_{12}q_{21}p_{11} + 2q_{21}p_{22}q_{22}q_{11}^2 q_{12} \\
&- q_{11}^2 p_{12}q_{12}q_{21}p_{21}q_{22} + q_{11}q_{21}^3 p_{22}q_{21}p_{21} + 2q_{11}p_{12}q_{12}q_{21}^2 p_{21}q_{22} - q_{11}q_{21}^3 p_{22}p_{11}q_{12} + 2q_{11}p_{12}q_{12}2_{21}^3 p_{21} \\
&\qquad - q_{21}^2 p_{22}q_{22}q_{11}p_{11}q_{12} + q_{21}^3 p_{22}q_{22}p_{21}q_{12}) \\
&\qquad \frac{(q_{12}p_{21}q_{22}^2)}{(p_{11}q_{11}q_{12}q_{21} + p_{11}q_{11}q_{12}q_{21} + p_{11}q_{11}q_{12}q_{21} + p_{11}q_{11}q_{12}q_{21})^3)}|_{q_{11}=q_{12},q_{21}=q_{22}} \\
&= a(-p_{12}q_{12}^2 q_{22}p_{21} + 2q_{22}p_{22}q_{12}^2 p_{11} + p_{12}q_{12}^3 p_{11} - q_{22}^2 p_{22}p_{11}q_{12} + 2p_{12}q_{12}q_{22}^2 p_{21} + q_{22}^3 p_{22}p_{21})p_{21} \\
&\qquad \frac{1}{4}\frac{1}{q_{12}(p_{11}q_{12} + p_{21}q_{22})^3},
\end{aligned} \tag{18}$$

where a is positive by construction. Furthermore, all prices and quantities are positive. Therefore, the fraction $\frac{1}{4}\frac{1}{q_{12}(p_{11}q_{12}+p_{21}q_{22})^3}$ is positive. To show that being richer gives a country a greater weight in the constructed reference price vector, it is sufficient to show that the following is positive:

$$-p_{12}q_{12}^2q_{22}p_{21}+2q_{22}p_{22}q_{12}^2p_{11}+p_{12}q_{12}^3p_{11}-q_{22}^2p_{22}p_{11}q_{12}+2p_{12}q_{12}q_{22}^2p_{21}+q_{22}^3p_{22}p_{21} \tag{19}$$

Applying maple to minimize expression (18) numerically, we find that this expression cannot be negative because its minimal value is zero. However, forcing all quantities and prices to be strictly positive causes the minimal value to be a small positive value. The maple output from the minimization procedure is given in Table 5.

The minimal value of the expression in (18) is zero when the prices of both goods in country 2 are zero. When both prices are zero in country 2, country 2's prices have no weight in the construction of the reference prices (from equation (16)). Therefore, the prices of country 1 are also given maximum weight when the countries are equally rich. Thus, becoming richer does not influence the weight.

However, the minimal value is never negative, and when forcing prices and quantities to be strictly positive, the effect of getting richer is positive. Hence, if prices and quantities are strictly positive, a richer country has a larger weight in the construction of the Geary–Khamis prices. Under special circumstances, a richer country is given an equal weight as a poorer country; however, the effect of being richer is never negative.

[Table 5 about here.]

The substitution effect is similar to the Gerschenkron effect in the growth literature, according to which, the earlier is the base year, the higher is the measured growth rate (Gerschenkron, 1947).

Appendix B Robustness Analysis

In this Appendix, four different tests of robustness of the estimates from the main model based on nine countries are conducted. The main results are confirmed in each of the tests.

First, the preferred specification given in (4) is estimated on the subset of households with two children and two adults to test whether differences in household composition influence the main results. Second, an alternative specification is considered by applying the OECD adult equivalence scale. Third, a semiparametric analysis is conducted to study whether the functional form fits the data used in the study. Fourth, relative prices are included and the standard AIDS specification given in equation (1) is estimated on the subgroup of our sample for which relative prices are available; i.e., on the households in the five countries in which cross-country comparable relative prices are available.

B.1 Household composition

The first robustness check is conducted by including all households rather than only a subset of households of same composition and size. This yields a much larger sample of households and hence more information for the analysis. However, an equivalence scaling is needed to make incomes comparable among households; we chose the OECD equivalence scale.[11] The regression results are reported in the first column of Table 2. The four non-OECD countries again have the highest bias according to these estimates, and the main result given in Figure 1 is confirmed (see Figure 7). Côte D'Ivoire has a smaller bias than in the preferred model and China has the largest bias followed closely by Nicaragua, Azerbaijan, and Côte D'Ivoire. Table 6 reports the Gini indices for this robustness check. The Gini index also increases substantially in this case, from 0.45 to 0.58. In addition, the finding of Lorenz dominance of the uncorrected measures is confirmed (see Figure 8). Therefore, analyzing only a subsample of households does not seem to be crucial for our results.

[Figure 7 about here.]

[11] The OECD adult equivalence scale gives a weight of unity to the first person in the household, 0.7 to each additional adult, and 0.5 to each additional child (less than 16 years of age). The number of households differs substantially between countries. Despite this, the weight given to each household is the same. Two different weighting techniques were used for the robustness analysis: we used a weight equal to the population in the respective household's country and a weight equal to the ratio of observations relative to the population of the country of residence. Neither weighting scheme changes the main result.

[Table 6 about here.]

[Figure 8 about here.]

B.2 Functional form – does the semilog specification fit the data?

[Figure 9 about here.]

A major concern with the method applied in this paper is to what extent the functional form specification is restrictive. To study the functional form, a semiparametric analysis based on differencing is conducted. All variables, except the logarithm of real income, are included linearly in the regression. This robustness check, therefore, investigates whether the log-linear relationship between the budget share for food and real income fits the data well. Figure 9 shows the kernel regression between the budget share for food and the logarithm of real income after removing the effects of the other variables by differencing. The kernel regression function is linear where the curve is precisely defined; i.e., where the upper and lower bounds from the bootstrapping coincide with the kernel itself. The semiparametric analysis, therefore, confirms that the log-linear relationship between the budget share for food and real income assumed in equation (2) fits the data well. As expected, conclusions are more robust in medium to high income levels for which there are more observations than are available in the lower tail of the income distribution.

For the lower incomes, the line does not seem to be perfectly linear, and we have a group of observations to the southwest of the line of the specified functional form. If this is the true functional form, the fact that the curve is not perfectly linear would exert a negative effect on the dummies for the poorest countries. The estimated dummies for the poorest countries exceed unity; however, our estimates are conservative estimates if the functional form in Figure 13 is the correct one and is not log linear. However, we consider the line in Figure 13 to be log linear in the sense that we cannot reject the functional form assumed in this paper.

B.3 Including relative price effects

Given that comparable *national* relative prices are available for five of the nine countries in the study, we examine whether including these relative prices changes the main results of this paper. This is done by estimating equation (1) on the subsample of households in the five countries that have such prices available and by applying the significant relative price effect estimated by Costa (2001).

Given that the coefficient for the relative price has already been estimated by Costa (2001), one new *net* variable is constructed. The dependent variable is the difference between the budget share for food and the effect of relative prices given in equation (19). The coefficient of the logarithm of relative prices is assumed to be 0.006, which is Costa's estimated coefficient.[12] The new left-hand side variable is defined as:

$$m^c_{h,j} = m_{h,j} - 0.006 * (ln(P_{f,j}) - ln(P_{n,j})). \tag{20}$$

Based on this variable, a new regression is run, and a new set of dummy coefficients and subsequently a new set of PPP biases are estimated. The estimation results are given in the third column of Table 2.

Again, in this case the PPP bias is a function of the coefficient of the logarithm of real income and the country dummy coefficient. In addition, it is now also a function of the bias in the measured prices for food and nonfood items, and is given by:

$$\ln E_j = \frac{\gamma}{b}(\ln E_{f,j} - \ln E_{n,j}) - \frac{d_j}{b}, \tag{21}$$

where $E_{f,j}$ and $E_{n,j}$ are the biases in the measured prices for food and nonfood items, respectively. Because we have no method for identifying the bias in all three prices si-

[12] The corresponding price elasticity is approximately 0.68. This price elasticity is calculated as $-1 + [(\gamma - \alpha b)/m]$, where α is the share of food in the United States total price index (Costa, 2001).

multaneously, we assume that the bias in the price for food cancels out the bias in the nonfood price; i.e., we assume that there is no bias in the *relative* price. Under this assumption, the bias is measured as expressed in equations (3) and (6). This assumption is quite strong and because we cannot identify all the biases, we cannot test whether this assumption is valid. However, we know that the estimated coefficient of relative prices is well below the coefficient of real income and, therefore, the major effect picked up by the country dummy coefficient is from the PPP bias. Despite this, it should be kept in mind that if the bias for the food price is larger than the bias for the nonfood price, the bias is overestimated. The opposite applies if the bias for the nonfood price is larger than the bias for the food price.

For the five countries under study, the measured PPP bias is higher the lower is real income (see Figure 10). In addition, in this case, there is a negative relationship between the PPP bias and real income. The third and fourth columns of Table 6 report the Gini indices before and after the correction, respectively. We obtain a Gini index for these countries of 0.17 when using data from the PWT whereas, when correcting for the measured PPP biases, we obtain a substantially higher Gini index of around 0.30. As Figure 11 shows, the distribution of EC incomes Lorenz dominates the distribution of PWT incomes. Therefore, all four of our major findings are substantiated when relative prices are included in the analysis.

[Figure 10 about here.]

[Figure 11 about here.]

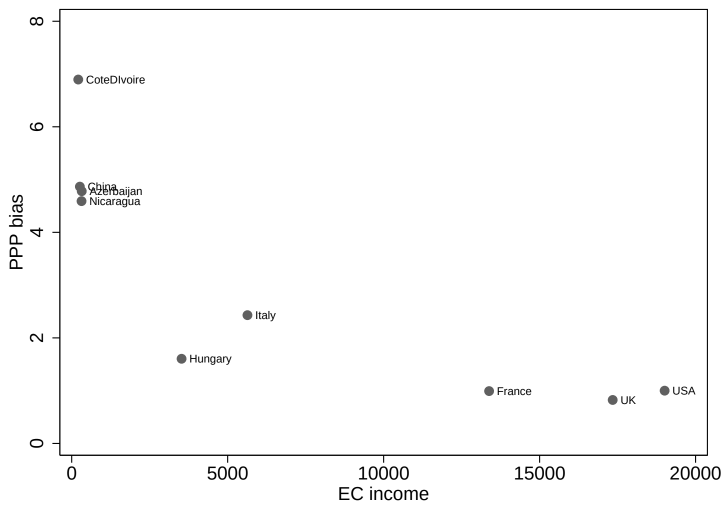

Figure 1: **PPP bias and real income.** The figure illustrates the relationship between the PPP bias and EC income for the nine benchmark countries. The bias for country j is the conversion factor for the PWT incomes, and the EC incomes are measured by the regression coefficients: $PPPbias_j = e^{\frac{-d_j}{b}}$.

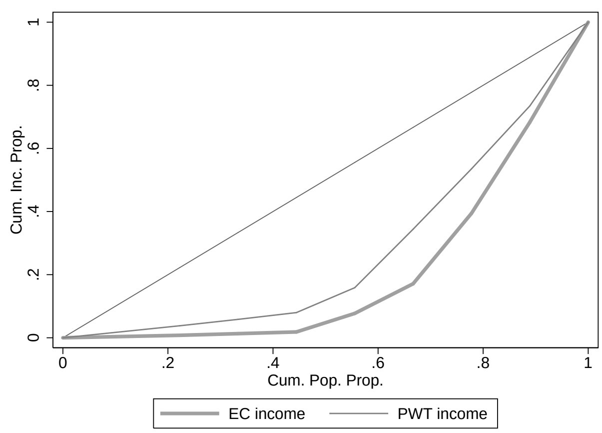

Figure 2: **Lorenz curves from the preferred model, EC incomes, and PWT incomes.**

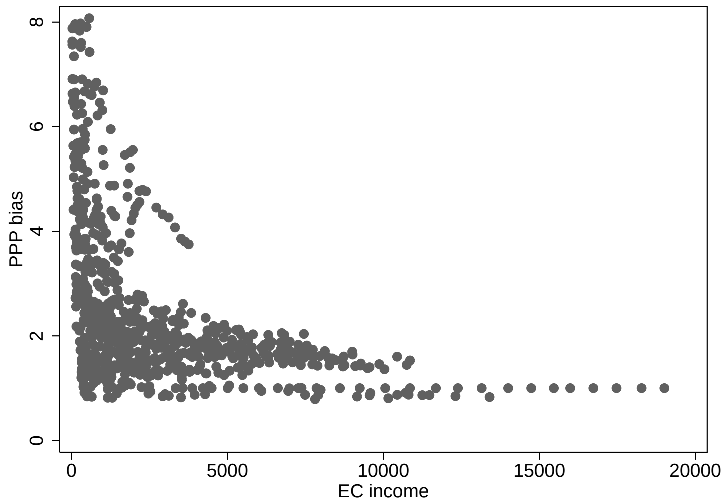

Figure 3: **The relationship between the PPP bias and real income – extended model.** The figure illustrates the relationship between PPP bias and EC income based on the 983 observations used for the extended model.

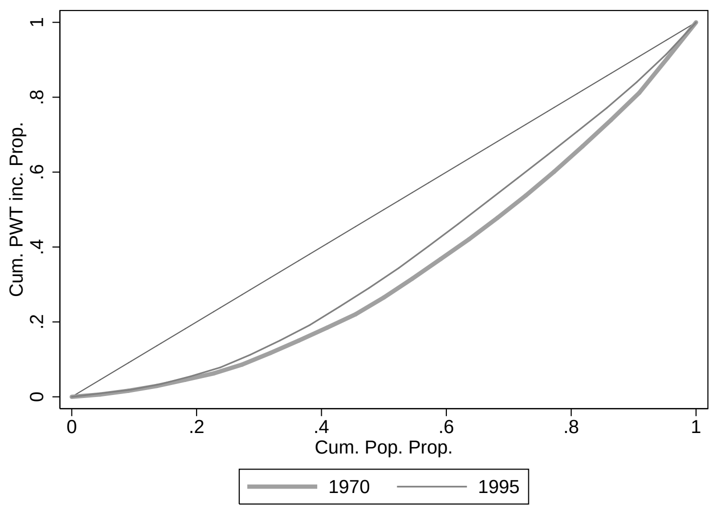

Figure 4: **The Lorenz curves for 1970 and 1995 based on PWT incomes.**

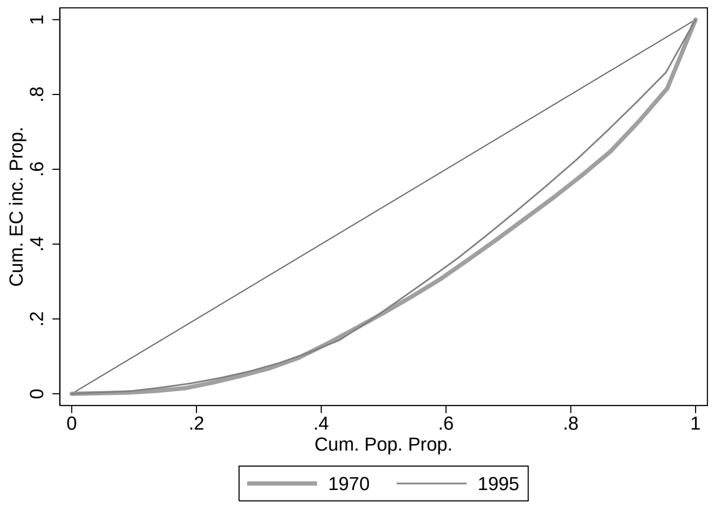

Figure 5: **The Lorenz curves for 1970 and 1995 based on EC incomes.**

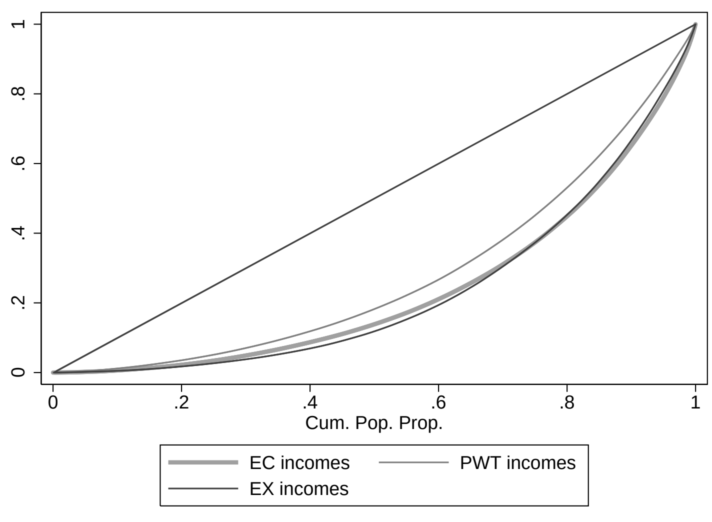

Figure 6: **The Lorenz curves for the EC, EX and PWT incomes.**

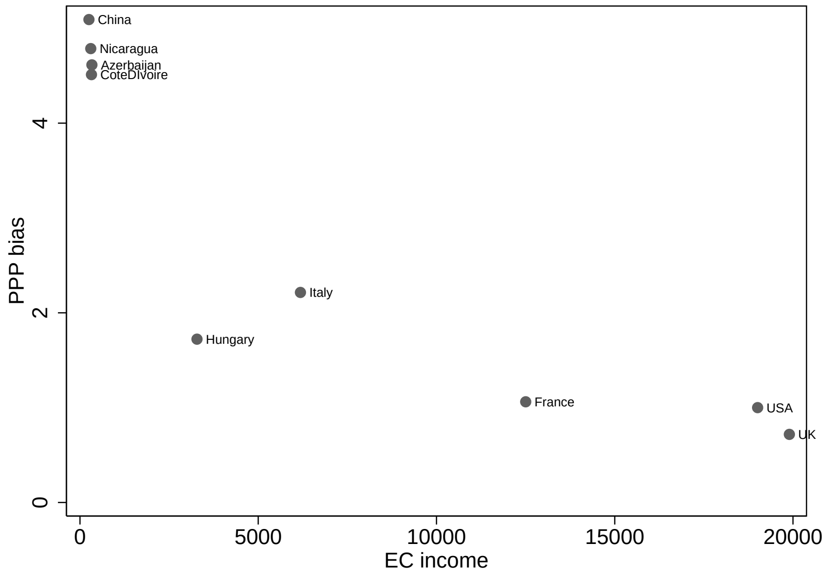

Figure 7: **PPP bias and per capita real income:** The figure illustrates the relationship between the PPP bias and EC income for the nine benchmark countries based on OECD adult equivalence scaling. The bias for country j is the conversion factor between the PWT incomes and the EC incomes based on the regression coefficients, as follows: $PPPbias_j = e^{\frac{-d_j}{b}}$.

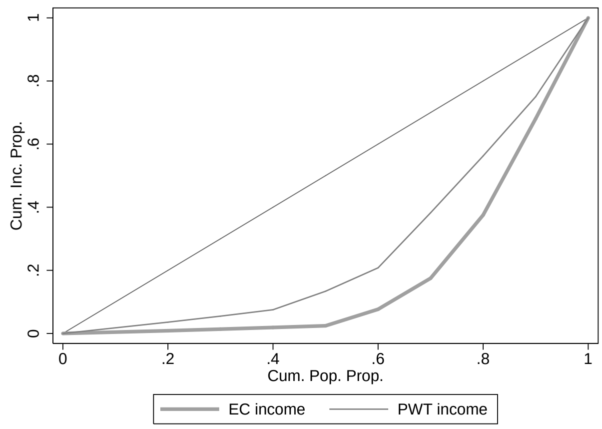

Figure 8: **Lorenz curves based on OECD adult equivalence scaling, EC incomes, and PWT incomes.**

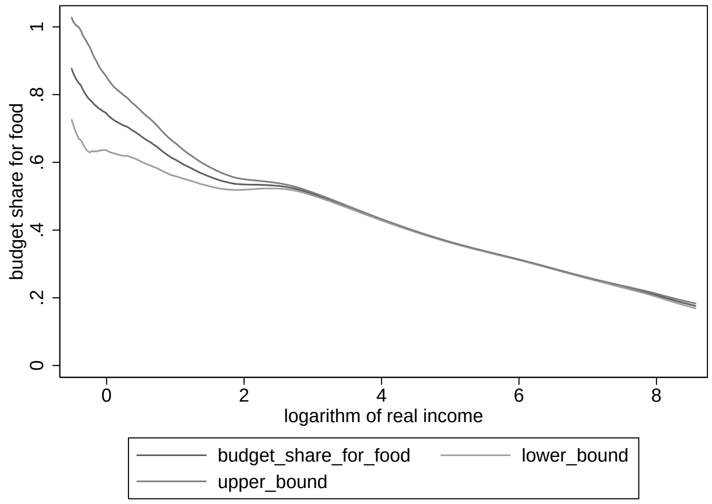

Figure 9: **Kernel regression.** The Kernel relation using the Epanechnikov kernel smoother: the relationship between the budget share for food and the logarithm of real income when the effects of the other explanatory variables are removed by differencing. Tenth-order differencing is conducted based on the optimal differencing weights proposed in Yatchew (2003). The bandwidth is equal to 1.36149. The bandwidth used is obtained from the formula $bandwidth = 0.15 * (max_{logrealcons} - min_{logrealcons})$, where $max_{logrealcons}$ and $min_{logrealcons}$ are the maximum and minimum values of the real logarithm of expenditure, respectively. The bounds correspond to the 95% confidence intervals. For values of the logarithm of real income below 2.7 the curve is not precisely defined. This is because of the small number of observations. It is not clear whether the curve is linear in this segment. However, to the right of this point, the curve is precisely defined and is linear.

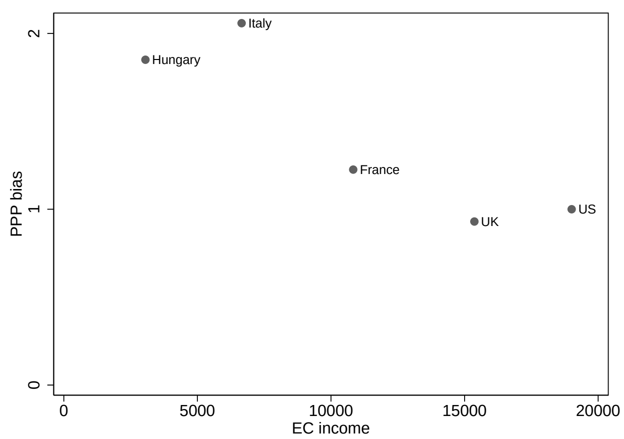

Figure 10: **PPP bias and per-capita real income:** The figure shows the relationship between the PPP bias and EC income.

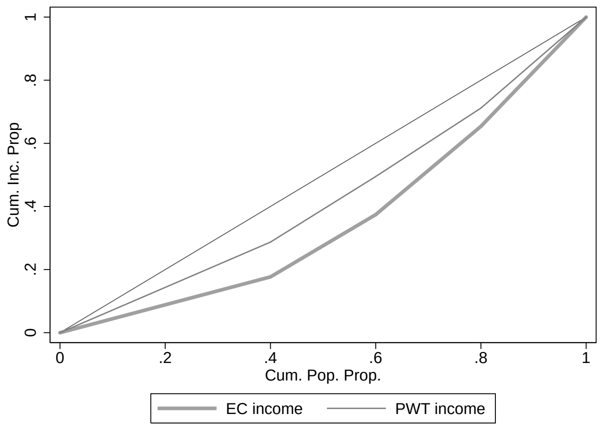

Figure 11: **Lorenz curves based on including relative price effects, EC incomes, and PWT incomes.**

	Survey year	Institution	No. of hh	Nat. Repr.
Azerbaijan	1995	SORGU / World Bank	1966	Yes
China	1994	Min. of Agg./World Bank	786	No
Côte D'Ivoire	1986	Inst. Nat. Stat. / World Bank	1573	Yes
France	1995	Inst. Nat. Stat. Étud. Éc. / LIS	9627	Yes
Hungary	1996	Hungarian Cent. Stat. Off.	7528	Yes
Italy	1995	Bank of Italy / LIS	8116	Yes
Nicaragua	1993	INEC / World Bank	3185	Yes
United Kingdom	1995	UK Data Archive / LIS	6789	Yes
United States	1995	CES, US Bureau of Labor	12973	Yes

Table 1: **The different surveys.** This table provides an overview of the nine different surveys included in the study and the institutions that conducted the surveys.

	Adeq	Same-Composition	Costa
log of real income (adeq adj.)	-.094		
	(.001)		
log of real income		-.097	-.089
		(.003)	(.0009)
Azerbaijan	.144	.151	
	(.005)	(.018)	
China	.153	.153	
	(.005)	(.012)	
Nicaragua	.147	.147	
	(.004)	(.011)	
Côte D'Ivoire	.142	.186	
	(.005)	(.019)	
Hungary	.051	.046	.055
	(.002)	(.007)	(.002)
France	.006	-.0008	.018
	(.001)	(.004)	(.001)
United Kingdom	-.031	-.019	-.006
	(.001)	(.004)	(.001)
Italy	.075	.086	.064
	(.002)	(.005)	(.002)
children	-.0002		.021
	(.0005)		(.0006)
adults	-.005		.029
	(.0005)		(.0006)
age	.0006	.001	.0006
	(.00003)	(.0002)	(.00003)
cons	.663	.727	.621
	(.006)	(.020)	(.006)
Number of observations	52543	4968	45033
Adjusted R-squared	.564	.619	.383

Table 2: **Regression results.** The table reports three sets of estimates (standard errors are in parenthesis). The estimates for the main model, represented by equation (2), and for the preferred specification, which is estimated for households of the same size, are reported in the second column. To avoid the inaccuracies generated by household composition, we analyze the subset of households with two children and two adults. The first column reports the estimates for equation (2), which is estimated for all households by using the OECD's adult equivalence scaling. The second column reports the coefficients from the preferred estimation for households with two children and two adults. The third column reports the estimates from a model that includes relative prices. The specification of this model is obtained from equation (1) by using Costa's (2001) estimated relative price coefficient, as specified in (19).

	Gini PWT	Gini EC
Benchmark countries	0.45	0.58
Pop w, benchmark countries	0.58	0.72
Extended model	0.37	0.46
Pop w, extended model	0.56	0.70

Table 3: **Gini indices.** The table shows the Gini index, as measured by the PWT incomes and the EC incomes. The first row presents the unweighted Gini index; i.e., the index that gives equal weight to each country irrespective of its size. The second row presents the population weighted Gini index, which weights each country proportionally to its population size. The third and fourth rows present the unweighted and weighted Gini indices, respectively, from the extended model.

Measure	1975	1995	**C**onvergence / **D**ivergence
PWT	0.32	0.25	C
EC real income	0.42	0.38	C

Table 4: **The Gini coefficient based on different measures, 1970 and 1995.**

Restrictions	q's ≥ 0 and p's≥ 0	q's >0 and p's> 0	q's≥ 1 and p's≥ 1
Minimal value	0	$0.114e-19$	$0.107e-9$
p11	2.596	3.904	1742.8
p12	0	$0.100e-8$	1.000
q11	1.770	2.400	837.4
q22	1.770	2.400	837.4
p21	0.847	$0.100e-8$	1.00
p22	0	$0.100e-8$	1.00
q21	1.029	1.015	37.15
q22	1.029	1.015	37.15

Table 5: **The maple output of the minimization.** The table shows the minimal value of (18) and the values for prices and quantities at the minimal value for three different restrictions on prices and quantities.

	Gini PWT	Gini EC	Gini PWT (Rel prices)	Gini EC (Rel prices)
Unweighted	0.45	0.58	0.17	0.30
Pop weighted	0.58	0.73	0.06	0.14

Table 6: **The Gini indices.** The table shows the Gini indices for the PWT incomes and the EC incomes based on estimation for all households, and adjusting for the OECD's adult equivalence scaling, and for the EC incomes based on including relative prices for the subsample of households from the OECD countries and based on applying Costa's coefficient for relative prices.