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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 incomes. The bias is substantial and systematic: the poorer 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 subsequent corrected incomes are measured by estimating Engel curves for food, which is an established method of measuring CPI bias.

JEL-Code: D10, E31, F01.

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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 policy makers. The magnitude of the differences, however, depends on the measure used for comparisons. To illustrate, (per capita) income in China is more than five times larger if one uses Penn World Table (PWT)[1] incomes rather than exchange rate based (EX) incomes.

In this paper, we study PWT incomes, which aim at correcting for price level differences across countries, and identify the bias in them by estimating Engel curves for food.[2] Furthermore, the relationship between the bias and the income of a country is studied. The PWT produces purchasing power parity (PPP) adjusted incomes, and thus the associated bias is referred to as the PPP bias. Having estimated the bias in PWT incomes, we provide new estimates of (real) income and refer to these as the Engel Curve (EC) incomes. By comparing the estimated EC incomes and the PWT incomes, the issue of how the bias influences estimated inequality is discussed. Finally, we discuss whether EX incomes, which simply transform each country's nominal income into one common currency, provide better estimates of income than do PWT incomes.

This paper reports three main findings. First, there is substantial and systematic PPP bias in the PWT incomes; the poorer the country, the more its income tends to be overestimated. Second, the PPP bias causes a substantial and robust underestimation of international inequality; the Gini index increases substantially when one adjusts for the bias. Third, whereas PWT incomes provide better estimates than the EX incomes for the richer countries, the EX incomes, which implicitly assume that PPP holds, provide better estimates for the poorer countries.

As we know that price levels differ across countries, there is consensus that the seminal work on establishing the PWT was a well-founded initiative, and the data have been extensively used.[3] Still, although many studies rely on PWT data, few focus on the PPP

[1] Heston et al. (2002).

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

[3] The early work on the PWT was conducted at the University of Pennsylvania by Irving Kravis, Alan Heston, and Robert Summers.

bias in this data set. Some contributors focus on one component of the bias, however, the so-called substitution bias, and use macro data to measure this bias (Dowrick and Akmal, 2005; Hill, 2000; Neary, 2004; Nuxoll, 1994). In these studies, it is shown that international income differences tend to be underestimated by the PWT data. However, because they only study the substitution 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 the calculation of bias corrected incomes, i.e. the 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 the method of Hamilton (2001) for estimating CPI bias to the estimation of the PPP bias.[4]

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. Since Ernst Engel's work (1857;1895) we have had the notion of an empirical regularity: As income increases, the budget share for food decreases. As Houthakker (1987) states, of all empirical regularities observed in economic data, Engel's law is probably the best established. We use this empirical regularity and make the assumption that is standard in the Hamilton tradision, namely that there is a stable relationship between the budget share for food and household 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 reference country, in our case the United Kingdom, is interpreted as PPP bias for that country relative to the United Kingdom.

The paper is organized as follows. In Section 2, we discuss the causes of the bias and

[4]We also extend the Hamilton method by fully incorporating the quadratic extension suggested by Costa (2001).

why the PWT tends to be systematically biased. In Section 3, we describe the empirical methodology in detail. In Section 4, we describe data used in the main analysis. The analysis and main findings are presented in Section 5. Section 6 contains the robustness analysis. Section 7 extends the analysis by using UN aggregate consumption data, and Section 8 evaluates EX incomes and compares this evaluation to that of the PWT incomes. Section 9 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 any observed price difference for cars between Poland and the United States reflects a difference in 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 in practice 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. Hence, unless the quality differences are fully adjusted for, both PPP and CPI measures incorporate a quality bias.

Second, the substitution bias arises because a reference price vector is applied to evaluate different countries' *realized* consumption bundles. The fact that the consumers, unless they have Leontief preferences, would substitute 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.[5] Hence, unless consumers have Leontief preferences, both PPP and CPI measures incorporate substitution bias.

Both the quality bias and the substitution bias are expected to be systematic. Because we may expect that poorer countries have products of lower quality than richer countries, it follows straightforwardly that failing to adjust for quality causes poorer countries' incomes to be overestimated. Interestingly, we also expect the substitution bias to cause an overestimation of poorer countries' incomes relative to richer countries' income. Independent of income level, the substitution bias always leads to an overestimation of a country's income. This overestimation is larger the larger the difference between the own price vector and the reference price vector (Nuxoll, 1994). The Geary–Khamis reference prices are by construction closer to the prices of the countries with larger total income, and, hence, we expect the substitution bias to be larger for the countries with lower income than for those with larger income.

The left panel of Figure 1 shows the relationship between the weight of a specific country in the construction of the reference prices underlying the PWT, and the total income of this specific country. Country j's weight is defined by the difference between the Geary–Khamis reference prices when including all countries, and the reference prices when including all countries but country j.[6] A country's income is measured by the PWT. We can see that richer and larger countries influence the reference price level more as the weight in reference prices is increasing in the total income of a country. The solid line represents the fitted line from regressing the logarithm of the difference on the logarithm of per capita income; the coefficient being 0.906 (*p-value* < 0.001).

[Figure 1 about here.]

Not surprisingly, as shown in the right panel of Figure 1, we also identify a positive relationship between this weight and per capita income. The two solid lines represent the

[5]The Geary–Khamis price indices are Laspeyres indices as they compare each country's price level with the constructed price level.

[6]The difference for country j, d_j, between the two constructed price vectors is calculated by the following formula: $d_j = \frac{\sqrt{(\Sigma_i^{11}(x_i - y_{ij})^2)}}{\Sigma_i^{11} x_i^2}$ where x_i is the reference price of good i when all countries are included in the construction and y_{ij} is the reference price of good i when all countries except country j are included in the construction.

fitted line from regressing the logarithm of the difference on the logarithm of per capita income; the upper line displays the result of this regression when weighting by population size (the coefficient being 0.84 (*p-value* < 0.000)) whereas the lower line shows the result of an unweighted regression (the coefficient being 0.420 (*p-value* $= 0.024$)). The countries in the middle of the per capita income distribution with very small weights are very small countries such as St. Kitts and Nevis, and Antigua and Barbuda.

3 Empirical Methodology

If households with the same PWT measured income and the same demographic characteristics systematically report a different budget share for food in one country than in another, we attribute this systematic difference 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 the level of household income, and, subsequently, to the PPP bias in this 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, we have evidence from studies of different countries and over different periods, 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).

In order to allow for some functional form flexibility, we estimate two demand systems. First, we follow Hamilton (2001) and estimate the Almost Ideal Demand System (AIDS) (Deaton and Muellbauer, 1980). Second, we estimate the quadratic extension of this system, the Quadratic Almost Ideal Demand System (QUAIDS) (Banks et al., 1997). Below, we present the two systems and show how the PPP bias is measured within each of them. The estimates, and subsequent results, from the two systems are very similar.

3.1 The Almost Ideal Demand System

The Engel curve of the AIDS is given by:

$$m_{h,r,j} = a + b(\ln y_{h,r,j} - \ln P_j) + \gamma(\ln P^f_{r,j} - \ln P^n_{r,j}) + \theta X_{h,r,j} + \varepsilon_{h,r,j}, \tag{1}$$

where $m_{h,r,j}$ is the budget share for food, $y_{h,r,j}$ is the nominal household income measured in 1996 United States dollars, and $X_{h,r,j}$ is a vector of demographic control variables including the age of the household head and the number 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 non-food items in region r in country j.

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

$$P_j = P'_j * E_j. \tag{2}$$

Equation (1) can therefore be expressed as:

$$m_{h,j} = a + b(\ln y_{h,j} - \ln P'_j) + \gamma(\ln P^f_{r,j} - \ln P^n_{r,j}) + \theta X_{h,j} + \sum_{j=1}^{N} d_j D_j + \varepsilon_{h,j}, \tag{3}$$

where D_j is the country dummy. The country dummy coefficient, d_j, is a function of the PPP bias, E_j, and the coefficient for the logarithm of household income, b:

$$d_j = -b \ln E_j. \tag{4}$$

From Equation (3) it follows that the PPP bias is given by:[7]

$$E_j = e^{-\frac{d_j}{b}}. \tag{5}$$

[7] Our main results are robust to measuring the PPP bias through the expenditure function of the demand system (see appendix B for details).

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

3.2 The Quadratic Almost Ideal Demand System

The Engel curve of the QUAIDS is given by:

$$m_{h,r,j} = a + b_1(\ln y_{h,r,j} - \ln P_j) + b_2(\ln y_{h,r,j} - \ln P_j)^2 + \gamma(\ln P^f_{r,j} - \ln P^n_{r,j}) + \theta X_{h,r,j} + \varepsilon_{h,r,j}. \tag{6}$$

Equation (6) can be expressed as:

$$m_{h,r,j} = a + b_1(\ln y_{h,r,j} - \ln P'_j - \sum_{j=1}^{N} d_j D_j) + b_2(\ln y_{h,r,j} - \ln P'_j - \sum_{j=1}^{N} d_j D_j)^2 + \gamma(\ln P^f_{r,j} - \ln P^n_{r,j}) + \theta X_{h,r,j} + \varepsilon_{h,r,j}. \tag{7}$$

where D_j is the country dummy, picking up the PPP bias directly. The country dummy coefficient is equal to the log of the bias

$$d_j = \ln E_j. \tag{8}$$

Consequently, for the QUAIDS, the PPP bias is given by:

$$E_j = e^{d_j}. \tag{9}$$

3.3 The different income measures

The relationship between EX, PWT, and EC incomes can be shown as follows:

$$Y_j^{EX} = Y_j \quad , \quad Y_j^{PWT} = \frac{Y_j}{P'_j} \quad , \quad Y_j^{EC} = \frac{Y_j}{P_j} = \frac{Y_j}{P'_j E_j}$$

where Y is the nominal per capita income in United States dollars. If the bias exceeds unity, the PWT consumption price is underestimated and, therefore, the income of the country is overestimated. The larger the estimated country dummy coefficient, the larger is the estimated bias, and consequently, the more the national per capita income is overestimated.

4 Data

We start out by using household micro data on ten base countries, one from each decile of the PWT income distribution, to estimate Engel curves for food. Table 1 provides an overview of the different surveys. The household data for Azerbaijan, Brazil, Bulgaria, Côte D'Ivoire, Nepal, Peru, and Tanzania are from the World Bank's living standard measurement surveys (LSMS).[8] The Hungarian data are from the Hungarian Central Statistical Office (Household Budget Survey Section). The Spanish data are provided by Instituto Nacional de Estadística (INE) and the data for the United Kingdom are taken from two different sources: the National Food Survey (National Statistics) provides the information needed to obtain regional food prices whereas the Family Expenditure Survey (ONS) provides household expenditure information.

The ten base countries all participated in the benchmark price survey for PWT 6.1. The base year for PWT 6.1 was 1996, and hence the household surveys included are conducted as close as possible to 1996.[9]

[Table 1 about here.]

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

Many of the households included in the sample are farm households, for which home-produced food accounts for much of the total household consumption. We account for

[8] Detailed information on different LSMS is provided on the World Bank website (World Bank, 2005).

[9] Given available data, we were unable to find any survey for a country in the third decile closer to 1996 than the Cte d'Ivoirian study.

this by incorporating estimated market value of home-produced goods in the expenditure variable.

One limiting criterium is that in order to include the relative price control in equations (3) and (7), the surveys need to have price information on food items. The ten surveys include information either on prices for food items at household level (Azerbaijan, Brazil, Bulgaria, Peru, Tanzania)[10], or on quantities of food items consumed which enabled us to calculate unit values (Côte d'Ivoire, Hungary, Nepal, Spain, UK). As is well-documented in the literature, one problem related to using unit values and prices for broad item groups reported at household level, is that they depend on both quality and price (Deaton (1987;1988), Nelson (1990), McKelvey (2010)). For example, if a unit value or a unit price of meat is recorded, a lower price for one household could indicate either that this household faces lower prices *or* that it consumes lower quality meat. In order to adjust for quality, we follow the approach in Deaton et al. (2004)[11]: the logarithm of the unit value of each good is regressed on a set of regional dummies, the logarithm of household consumption, and demographic controls. This estimated relationship is then used to calculate the regional mean log prices using the sample means for the logarithm of expenditure and the demographic controls. We do not have unit values or prices for all items in all countries, and hence we use the weighted country-product-dummy (WCPD) method due to Rao (1990;2005) to identify an overall price of food in the different regions in our sample.[12]

Whereas the food items have defined quantities and thus can be converted into the same units (kilograms) for all countries, the non-food item units are not standardized across the different micro data sets. As we are not able to trace non-food prices from the micro data, we deflate by the ICP non-food price which we find by applying the WCPD method on ICP data. This is not an ideal procedure, as we expect the ICP data to be biased, but it turns out that our main findings are robust to different ways of incorporating

[10]For Peru and Tanzania, the micro data contain a detailed price survey at cluster/district level, but in order to aggregate up to item groups comparable across countries, we used household specific consumption weights, and hence the item prices we have are household specific.

[11]This is a modified version of Coondoo et al. (2004).

[12]As explained in Diewert (2005), in the case of two countries, the logarithm of the WCPD index provides a second order local approximation to the Törnqvist index.

relative prices. Appendix A discusses the calculation of relative prices in detail and shows robustness analysis related to the relative price inclusion.

The macro price variable, P'_j, is a composite price index for all consumption goods in country j, which is constructed using the Geary–Khamis method. The macro price variable for consumption and the exchange rate are taken from Penn World Table 6.1 (Heston et al., 2002). The consumption price in the PWT is reported in current prices, with 1996 United States dollars as base, and we use the United States exchange rate and CPI to make income levels comparable across countries and time. The United States CPI is taken from the World Bank's World Development Indicators online to calculate nominal incomes which are comparable across time (World Bank, 2007).

5 Analysis and Findings

In this section, the PPP bias is estimated by using household surveys from the ten countries, and the findings are discussed in detail.

[Table 2 about here.]

The regression results are presented in Table 2. The estimated income elasticity for food is in line with previous studies (Costa, 2001; Hamilton, 2001; Beatty and Larsen, 2005; de Carvalho Filho and Chamon, 2006). By construction, the United Kingdom 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 Kingdom. All countries have a positive dummy coefficient; i.e., the macro price variables in the PWT underestimate the macro price levels relative to the United Kingdom macro price level. Therefore, according to the EC method, all countries' incomes are overestimated relative to the income of the United Kingdom.

[Figure 2 about here.]

Figure 2 reports the relationship between the PPP bias (resulting from the estimates in columns one and two of Table 2) and income. This relationship reveals the first main finding: there is a negative relationship between the PPP bias and income. This is in line with the theoretical discussion of Section 2. As expected, we find that the poorer a country, the larger the PPP bias.

Table 3 shows the measured PWT, EC, and EX incomes for the ten base countries. We can see that for the countries in the six poorest deciles, Tanzania, Nepal, Côte d'Ivoire, Azerbaijan, Peru, and Bulgaria, the EC income is substantially closer to the EX income than to the PWT income. Spain has an EC income that is closer to the PWT income than to the EX income, whereas the middle income countries, Hungary and Brazil, have an EC income with approximately equal distance to the EX income and the PWT income.

[Table 3 about here.]

Table 4 reports our second main finding, which is that international inequality is substantially underestimated. The table shows that the Gini index increases substantially when adjusting for the PPP bias; the first row shows that the unweighted Gini index increases from 0.50 to 0.64 for the base countries when adjusting for the bias, and the second row shows that the population-weighted Gini index increases from 0.39 to 0.48.[13]

[Table 4 about here.]

6 Robustness Analysis

In this section, we provide several robustness checks that all confirm the main results. First, the specifications given in equations (3) and (7) are estimated using all households independent of size and composition. Second, the fit of the two demand systems is discussed and a semiparametric analysis conducted. Third, we replace the Engel curve for food with an Engel curve for calories.

6.1 Household composition

The first robustness check is conducted by including all households rather than only a subset of households of same composition and size. The regression results are reported in the

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

third and fourth columns of Table 2. Again, we find a negative relationship between PPP bias and income and our main results are confirmed (see Figure 3, first row). Therefore, using only the subsample of households with two children and two adults is not crucial for our results.

[Figure 3 about here.]

6.2 Functional form

To test the robustness of the functional form assumptions, we have estimated two demand systems, the AIDS and the QUAIDS, which allows for some flexibility. We can see in Table 2 that the two systems give very similar results, which indicates that the choice of either one of the systems is not crucial to our results. We can see that the coefficient for the square of the logarithm of income is insignificant in our preferred estimation where we only include households with two children and two adults and hence, for this sample, we are unable to reject a hypothesis stating that the budget share for food is log-linearly related to the budget share for food (see e.g., Banks et al. (1997) for the same finding). However, when including all households in the estimation, the coefficient becomes significant.

To look more closely at the functional form assumption, we present a semi-parametric analysis. Figure 4 shows the kernel regression displaying the Engel relationship between the budget share for food and the logarithm of income after removing the effects of the demographic variables by differencing. We can see that it is very close to log-linear. However, in the lower tail of the income distribution where we have fewer observations, the bounds are wider and we cannot determine with precision the functional form in this area.

In sum, the empirical analysis confirms that we have no reason to expect that the functional form assumptions drive the results of this paper.

[Figure 4 about here.]

6.3 Engel curves based on calories

Food is a composite good and it might be the case that richer households consume higher quality calories, such as those from eggs and meats, whereas poorer households consume lower quality calories, such as those from wheat and rice. If this is the case, our estimated Engel curve is potentially a composite of calories and food quality. In this section we suggest replacing the Engel curve for food with an Engel curve for calories. We estimate the calorie content of the food basked for all households in our sample with two children and two adults by using calorie tables (Nutribase, 2001). Hence, we can calculate the household-specific price of calories as:

$$p_h^c = \frac{exp_h^f}{cal_h}, \tag{10}$$

where exp_h^f is total expenditure on food and cal_h is number of calories consumed by household h.

We know that the household-specific price of calories is a function of the price of food items that the household faces, but potentially also a function of the quality of the food that the household consumes. In order to trace the quality adjusted budget share for calories, we need to find a quality-adjusted price. Hence, we proceed to find the relationship between p^c and income and demographics by estimating the relationship between p^c and log of income and demographics, including regional fixed effects. Under the assumption that the price of different food items is regional specific, we find the quality adjusted calorie price, p^{cq}, by inserting the mean income into this relationship. This quality adjusted price is in turn used to calculate the budget share for calories as follows:

$$m_h^c = \frac{p_h^{cq} * cal_h}{p_h^{cq} * cal_h + exp_h^n}, \tag{11}$$

where exp_h^n is household h's total expenditure on non-food items. We then estimate the Engel curves given in equations (3) and (7) by using the budget share for calories from Equation (11) as the left hand side variable and the relative (quality adjusted) price of calories as a control in addition to the demographic controls. The estimation results are given in columns five and six in Table 2 and the subsequent relationship between PPP bias and income is provided in second row of Figure 3. We observe that the overall picture is very similar to that of the main analysis: The poorer a country, the larger the PPP bias.

7 An extended analysis

It is well known that micro data from household surveys and aggregate data may give quite different measures of income (see e.g., Deaton (2005)). In order to study whether the national data would reveal a different PPP bias than the survey data, we provide an extended analysis based on UN national mean variables. The extended analysis uses the estimated coefficient from the analysis on the ten base countries and UN mean variables (UN, 2008). Given a country's budget share for food and mean demographic characteristics, we attribute any difference between the PWT income and the EC income, to PPP bias. From equation (1) and aggregation to per household mean budget shares (see e.g. Denton and Mountain (2004)), it follows that:[14]

$$\overline{m_j} = a + b\frac{\overline{\frac{y_j}{P_j}\ln(\frac{y_j}{P_j})}}{\overline{\frac{y_j}{P_j}}} + \theta\overline{X_j}, \tag{12}$$

where $\overline{V}$ indicates the mean value of any variable V. The mean household demographic characteristics consist of predicted mean age of the household head, mean number of adults and mean number of children in the households.

We have estimated the coefficients for this model based on the micro data, and hence we can identify the term $\kappa = (\overline{\frac{y_j}{P_j}\ln(\frac{y_j}{P_j})})/(\overline{\frac{y_j}{P_j}})$ as follows:

$$\widehat{\kappa} = \frac{\overline{m_j} - \widehat{a} - \widehat{\theta}\overline{X_j}}{\widehat{b}} \tag{13}$$

where $\widehat{a}$, $\widehat{b}$, and $\widehat{\theta}$ are the coefficients estimated in our base model based on the ten countries for which we have micro data.

[14]Note that we do not have information on relative prices for the countries in the extended analysis. Hence, we implicitly assume that relative prices are the same across countries in this part of the analysis.

The PPP bias for country j is measured indirectly by using the estimated term $\widehat{\kappa}$, and the same term measured by using PWT prices ($\kappa' = (\overline{\frac{y_j}{P'_j}\ln(\frac{y_j}{P'_j})}/\overline{\frac{y_j}{P'_j}})$) (cf. Equation (2)):[15]

$$E_j = exp(\frac{\widehat{\kappa}}{\kappa'}) = \frac{\widehat{P}_j}{P'_j} \tag{14}$$

7.1 Data used in the extended analysis

We extend our analysis by using aggregate household data from the UN Statistics Division (Common Database). We include 32 observations on mean household consumption and budget shares, covering 32 countries in the year 1995. We use data on final household expenditure in national currencies at current prices.[16] To make final household consumption comparable across countries, we use the PWT price of consumption and the PWT exchange rate (Heston et al., 2002).

To simulate the distribution of consumption within each country, we assume that income is lognormally distributed. We use standard deviation for each country calculated by using the distributions estimated by Sala-i-Martin (2006). From the simulated distributions, we then calculate $(\overline{\frac{y_k}{P'_k}\ln\frac{y_k}{P'_k}})/(\overline{\frac{y_k}{P'_k}})$.

Information on demographic controls is also obtained from the UN (UN, 2008): The number of children and adults, and subsequently the OECD's adult equivalence scaling, can be calculated directly (UN Statistics Division, series codes 13681 and 1070). The age of the household head is predicted from observations on mean age of male citizens (UN Statistics Division, series code 13630) combined with the estimated difference between the mean age of household head in nine micro data sets and mean age of male citizens from the UN for the same nine countries (difference between them equal to 5.93). Hence, we predict mean age of household head by adding 5.93 years to the UN observations on

[15]As $\overline{\frac{y_j}{P_j}\ln(\frac{y_j}{P_j})}$ is generally different from $\overline{\frac{y_j}{P_j}}\ln(\overline{\frac{y_j}{P_j}})$, the former is simulated by using distributions from Sala-i-Martin (2006) and the assumption of lognormal distribution of income (see also Section 7.1).

[16]We use Table 3.2 in the UN statistics division, Common Database, and include all series in the 1993 SNA, i.e. series 100, 200, 300 and 400, where we have data on mean age of adult male population, mean household number of children and adults. We have to drop Azerbaijan and Namibia, however, the former because the final household consumption excludes some direct purchases and the latter because there is discrepancy between the components of consumption and final household consumption.

mean age of male citizens.[17]

PWT income is defined as the consumption level, measured by the consumption share of real gross domestic product per capita, whereas EX income is constructed by multiplying PWT income by the price of consumption , i.e., by eliminating the price deflation.[18]

7.2 Analysis and findings – extended analysis

We estimate the PPP bias for 32 countries in 1995.[19] As shown in Figure 5, also for this larger sample of countries we find that the poorer the country, the larger the bias. A more detailed description of the results is given in Table 5 which reports the EC income and the measured bias for the 32 countries.

[Figure 5 about here.]

[Table 5 about here.]

Table 4 shows that measured inequality for these countries increases substantially when the PPP bias is adjusted for, and hence our second main finding also carries through when using aggregate data.[20] Table 6 shows the estimation results from regressing the PPP bias against the log of income. We can see that the regression reports a strong negative relationship between the PPP bias and income (coefficient of -0.970) and that the regression has a fairly high explanatory power (R-squared equal to 0.59).

[Table 6 about here.]

[17] The nine countries being Azerbaijan, China, Côte d'Ivoire, Nicaragua, Hungary, Italy, France, the United Kingdom, and the United States.

[18] Our EX incomes are thus also very much dependent on the PWT (just not the price deflator of consumption). Other sources of exchange rate based incomes may differ from the exchange rate based incomes of this paper.

[19] N.N. (2008) includes more observations by introducing more years and hence duplicate income observations for many of the countries in the study. However, the results of this analysis are the same as the ones presented here, and hence, introducing duplicate observations for some countries does not add anything to the analysis.

[20] As we know that first, we are unable to control for relative prices, second, we use imputed distributions, and third, we work with aggregate data, we should be more focused on the systematic effect and pay less attention to the point estimates for each country.

8 Evaluating the EX Incomes

Historically, international comparisons of income have relied on the EX incomes, which transform incomes into a common currency, such as the United States dollar. Just as for the PWT incomes, we have reasons to expect that the EX incomes are biased. First, if either PPP does not hold, or if prices for nontraded goods differ between countries, then using the exchange rate yields biased estimates of income. Second, the quality bias would be equally important for the EX incomes as for the PWT incomes. We would also expect these two biases to be systematic, but systematic in different directions. As prices tend to be lower in poorer countries, it follows straightforwardly that failing to adjust for prices causes poorer countries' incomes to be *underestimated*. On the other hand, as we stated in Section 2, quality tends to be lower in poorer countries, and thus, failing to adjust for quality causes poorer countries' incomes to be *overestimated*.

Since we expect both the PWT incomes and the EX incomes to be biased, an interesting empirical issue is which approach provides the best estimates of income and, subsequently, international income inequality. This section discusses this question in two ways. First by comparing the estimated EC incomes and inequality measures of sections 5 and 7 to the EX incomes. Second, we apply the same method as for the PPP bias, and identify the EX bias through the EC method.

In Table 3 and 4, respectively, we compare the results from the different income measures. We observe that the EC income is closer to the EX income for the poorer countries, whereas the EC income is closer to PWT income for the richer countries. Table 4 shows that measures of international inequality based on the EC incomes are far closer to those based on EX incomes than to those based on PWT incomes.

The more direct way of identifying the EX bias is by estimating the AIDS and QUAIDS using nominal household income, $y_{h,j}$. The AIDS can be expressed as:

$$m_{h,j} = a + b' \ln y_{h,j} + \gamma(\ln P^{f}_{r,j} - \ln P^{n}_{r,j}) + \theta X_{h,j} + \sum_{j=1}^{N} d_j D_j + \varepsilon_{h,j}. \tag{15}$$

Subsequently, the EX bias is given by:

$$E_j^{EX} = e^{-\frac{d_j}{b'}}. \tag{16}$$

The QUAIDS can be expressed as:

$$m_{h,r,j} = a + b_1'(\ln y_{h,r,j} - \sum_{j=1}^{N} d_j D_j) + b_2'(\ln y_{h,r,j} - \sum_{j=1}^{N} d_j D_j)^2 + \gamma(\ln P_{r,j}^f - \ln P_{r,j}^n) + \theta X_{h,r,j} + \varepsilon_{h,r,j}, \tag{17}$$

where D_j is the country dummy, picking up the EX bias[21]. Consequently, the country dummy coefficient is equal to the log of the bias,

$$d_j = \ln E_j^{EX}, \tag{18}$$

and the EX bias is given by:

$$E_j^{EX} = e^{d_j}. \tag{19}$$

Table 2 (rows seven and eight) reports the results for these two estimations. Figure 6 shows the subsequent relationship between the EX bias and the EC income as well as that of the PPP bias and the EC income. The mean of the absolute bias for the base countries is equal to 0.23 for the EX incomes (0.93 for the extended analysis) and equal to 0.92 for PWT incomes (2.17 for the extended analysis).[22] This indicates that despite the empirical evidence against PPP, it is better to assume that PPP holds by using the EX incomes than to apply PWT incomes, when comparing incomes of both high- and low-income countries, e.g., when studying international income inequality.

When studying subgroups of countries at different income levels, however, this conclusion is relaxed. Dividing the base countries into two groups consisting of OECD and non-OECD countries, respectively, gives a mean of absolute bias for the OECD countries of less than 0.10 for the EX incomes (0.68 for the countries in the extended analysis) and

[21] Analogously to the PPP bias, the EX bias is defined as the factor that converts EC income into EX income.

[22] The mean of the absolute biases is calculated as $mean(|(bias-1)|)$. Hungary became a member of the OECD in 1996; for consistency we consider Hungary as non-OECD in both the base and the extended analysis.

less than 0.01 for the PWT incomes (0.44 for countries in the extended analysis). For the non-OECD countries it gives a mean of absolute bias of 0.33 for the EX incomes (1.40 for the countries in the extended analysis) and 1.71 for the PWT incomes (5.48 for the countries in the extended analysis). Hence, according to the EC incomes the measurement error for both PWT and EX incomes is larger for non-OECD countries than for OECD countries, and, moreover, the PWT incomes do better than the EX incomes for the richer countries, whereas the EX incomes do better than the PWT incomes for poorer countries.

[Figure 6 about here.]

9 Concluding Remarks

In this paper, we use household surveys from ten countries and UN mean household data to provide initial estimates of the overall purchasing power parity (PPP) bias in the Penn World Table (PWT). Although the PWT incomes are extensively used by economists, there are few studies investigating the bias in these measures. We find evidence of a substantial and systematic bias, and provide an 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 substantial and robust underestimation of international inequality. However, if studying a subgroup of richer countries only, the PWT seems to give more precise income estimates than the EX method.

The PPP bias is so substantial that applying the EX incomes, which implicitly assume both that PPP holds *and* that prices for nontraded goods do not differ across countries, yields better estimates of international inequality. However, if we concentrate on the OECD countries, PWT incomes give better estimates than EX incomes, whereas if we concentrate on non-OECD countries, EX incomes give more precise estimates than PWT incomes.

Several robustness checks show that the main findings are not driven by the misspecification of functional form, household composition, or quality effects.

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Appendix A Relative price of food and non-food

Based on the available food prices, we find it useful to harmonize the food categories by defining thirteen basic headings for food consumption: greens, meat, fish, salt, sugar, milk, egg, cheese, cereals, rice, soda, coffee, and oil. Whereas some of these item groups are rather small and probably contain quite comparable items across regions and countries, others are quite broad and there is a risk that they include different quality items. Hence, we need to adjust the prices for potential quality effects. Furthermore, there are different item categories recorded in the different countries and in order to arrive at the comparable item groups, we go through several steps.

First, we aggregate up to the thirteen basic headings at household level for the countries and items needed. For example the category *fish* consisted of different kinds of fish in some of the countries (and aggregation was needed), and in other countries there was reported one item called *fish*. The aggregation is done by using a Stone index weighting each price by the household budget share. Consequently, even with equal prices, a household consuming low quality fish has a lower reported price of fish than a household consuming higher quality fish. Hence, a second step adjusting the basic heading prices for potential quality effects, is needed. In order to adjust for quality, we regress the logarithm of each of these prices on a set of regional dummies, the logarithm of household consumption, and demographic controls. The regression coefficients are then used to predict the regional quality adjusted basic heading prices using the whole sample means for logarithm of expenditure and the demographic controls. The third step involves aggregation from the basic headings to an overall food price index. As some of the countries lack information on some of the basic heading prices, we use WCPD which allows missing values for some goods in some regions.[23] The non-food price is identified through using 1996 ICP data and WCPD aggregation into a food and non-food price.

Given the available data, there are three ways of incorporating relative prices. First,

[23] Azerbaijan has no price information on fish, soda, and cereals; Brazil lacks price information on bread, milk, and soda; Côte d'Ivoire has no price information on bread, sugar, coffee, and milk; Hungary has no price information on salt; Nepal has no price information on bread and soda; Peru has no price information on salt; Spain has no price information on eggs and cereals; Tanzania has no price information on salt, sugar, coffee, and soda; the United Kingdom has no price information on salt and rice.

we can use the food prices from micro data and the ICP non-food prices as we do in our main analysis. The resulting relative prices are displayed in Figure 7 (row one). Second, we can use the ICP data, and the regional variation from the food prices from the micro data. That is, we normalize the food prices so that the country average is equal to the ICP food price. These ICP relative prices are displayed in Figure 7 (row two). Third, we can simply use the food price from micro data as a relative price measure, implicitly assuming either that there is no cross price effect *or* that the non-food price is the same in all countries (these prices are shown in Figure 7, row three).

[Figure 7 about here.]

Figure 8 shows that although the point estimates for the separate countries change somewhat when using the alternative relative prices as controls, the systematic effect is preserved and hence our results are robust to these alternative ways of incorporating relative prices.

[Figure 8 about here.]

Appendix B Robustness analysis: PPP bias from the expenditure function

This appendix shows that our main findings are robust to using the expenditure function of the demand systems to identify the PPP bias, see Figure 9.

[Figure 9 about here.]

	Survey year	Institution	No. of hh	Decile
United Kingdom	1996	ONS and National Statistics	6412	10
Spain	1998	INE	14739	9
Hungary	1996	Hungarian Cent. Stat. Off.	7531	8
Brazil	1996	IBGE/World Bank	4898	7
Bulgaria	1995	Gallup International / World Bank	1886	6
Peru	1994	Cuánto S.A. / World Bank	3614	5
Azerbaijan	1995	SORGU / World Bank	1929	4
Côte D'Ivoire	1987	Inst. Nat. Stat. / World Bank	2899	3
Nepal	1995	CBS / World Bank	3372	2
Tanzania	1993	Planning Commission (UDS) / World Bank	5176	1

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

	AIDS	QUAIDS	AIDS ws	QUAIDS ws	AIDS cal	QUAIDS cal	AIDS ex	QUAIDS ex
Log of income	-0.105	-0.118	-0.100	-0.155	-0.122	-0.197	-0.105	-0.126
	(0.003)	(0.019)	(0.001)	(0.006)	(0.004)	(0.024)	(0.003)	(0.030)
Log of income sq		0.001		0.003		0.005		0.001
		(0.001)		(0.000)		(0.001)		(0.001)
Azerbaijan	0.075	2.067	0.120	4.019	0.143	3.771	-0.109	0.360
	(0.023)	(0.394)	(0.008)	(0.283)	(0.023)	(0.612)	(0.025)	(0.069)
Brazil	0.023	1.278	0.032	1.554	0.069	2.035	-0.018	0.859
	(0.006)	(0.102)	(0.002)	(0.051)	(0.012)	(0.218)	(0.007)	(0.068)
Bulgaria	0.112	2.989	0.135	4.454	0.119	3.158	0.006	1.092
	(0.010)	(0.354)	(0.004)	(0.199)	(0.014)	(0.446)	(0.012)	(0.129)
Côte d'Ivoire	0.124	3.300	0.164	6.432	0.162	4.267	0.035	1.423
	(0.019)	(0.594)	(0.006)	(0.387)	(0.022)	(0.711)	(0.019)	(0.256)
Hungary	0.056	1.752	0.093	2.941	0.055	1.872	-0.019	0.862
	(0.007)	(0.169)	(0.002)	(0.098)	(0.008)	(0.216)	(0.008)	(0.083)
Nepal	0.145	4.000	0.166	5.635	0.138	3.404	-0.035	0.728
	(0.012)	(0.509)	(0.004)	(0.260)	(0.013)	(0.449)	(0.016)	(0.093)
Peru	0.134	3.636	0.145	4.894	0.141	3.600	0.072	2.018
	(0.010)	(0.400)	(0.003)	(0.204)	(0.012)	(0.461)	(0.011)	(0.222)
Spain	0.000	1.028	0.010	1.164	-0.030	0.880	-0.017	0.876
	(0.010)	(0.104)	(0.003)	(0.045)	(0.006)	(0.071)	(0.010)	(0.089)
Tanzania	0.148	4.096	0.190	7.434	0.173	4.432	0.021	1.227
	(0.011)	(0.492)	(0.004)	(0.350)	(0.012)	(0.568)	(0.013)	(0.147)
Log of rel. prices	0.041	0.041	0.015	0.008	0.000	0.000	0.041	0.041
	(0.012)	(0.012)	(0.004)	(0.004)	(0.007)	(0.005)	(0.012)	(0.012)
Age	0.000	0.000	-0.001	-0.000	0.000	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Children			0.009	0.009				
			(0.000)	(0.000)				
Adults			0.018	0.018				
			(0.001)	(0.001)				
Constant	1.231	1.280	1.189	1.388	1.393	1.669	1.717	1.843
	(0.031)	(0.077)	(0.010)	(0.024)	(0.073)	(0.108)	(0.045)	(0.188)
Adj. R-Square	0.573	0.572	0.518	0.519	0.503	0.504	0.573	0.572
Number of Obs	4987	4987	52454	52454	4818	4818	4987	4987

Table 2: **Regression results, least squares estimation.** The table reports eight sets of estimates (standard errors are in parenthesis). The first and second columns report the estimates for the households with two children and two adults. The third and fourth column report the estimates for the whole sample (including all households independent of composition and size). The fifth and sixth columns report the coefficients for the calorie based Engel curves. The seventh and eight columns report the estimates using the exchange rate to make income comparable across households in different countries. The estimates of the main model (columns one and two) are discussed in Section 5, whereas the estimates of the robustness checks of columns three, four, five, and six are discussed in Section 6. The estimates reported in the seventh and eight column are discussed in Section 8.

	Y^{PWT}	Y^{EC}	Y^{EX}
UK	15088	15088	15088
Spain	11935	11897	10162
Hungary	5651	3324	2780
Brazil	4818	3857	3235
Bulgaria	3027	1040	1106
Peru	2839	895	1575
Azerbaijan	1739	852	303
Côte D'Ivoire	1471	453	634
Nepal	829	210	151
Tanzania	372	91	111

Table 3: **Three different income measures.**The table shows the income measured by PWT, EC incomes, and EX incomes for the ten base countries.

	Gini PWT	Gini EC	Gini EX
Base countries			
Unweighted	0.50	0.64	0.64
Population-weighted	0.39	0.48	0.49
Extended model			
Unweighted	0.26	0.39	0.34
Population-weighted	0.22	0.32	0.32

Table 4: **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 results for the extended analysis.

Country	Y^{EC}	E	Standard Error of E
Botswana	168	14.49	0.305
Belarus	231	12.33	0.439
Estonia	340	11.98	0.292
Latvia	433	7.88	0.178
Dominican Republic	545	4.79	0.081
Iran	592	5.23	0.135
South Africa	851	5.41	0.099
Colombia	1163	3.25	0.058
Mexico	2281	2.33	0.031
Hungary	2945	1.94	0.026
Israel	4356	2.11	0.019
Portugal	5535	1.69	0.012
Greece	5626	1.72	0.012
Spain	6460	1.86	0.012
New Zealand	6755	1.71	0.010
Italy	6847	2.00	0.011
Japan	7159	1.96	0.008
Ireland	7933	1.38	0.006
Hong Kong	8770	1.91	0.004
Belgium	9129	1.23	0.003
Norway	9629	1.43	0.006
France	9722	1.37	0.005
Finland	10234	1.16	0.005
Australia	11513	1.32	0.001
Austria	11826	1.20	0.002
Switzerland	12016	1.17	0.001
Sweden	12329	1.11	0.003
Canada	12567	1.04	0.001
Denmark	12604	1.21	0.002
Germany	13008	1.09	0.001
United Kingdom	14291	1	0
United States	15541	1.22	0.004

Table 5: **EC income, PPP bias and standard deviation of bias.**The table displays the EC incomes, the PPP bias and the standard error of the PPP bias for the 32 countries included in the extended analysis for the year 1995. The estimates from the base model in parenthesis. *The estimate is for 1996.

	Dep var: E	p-value	R-squared	N*
Log of EC income	-0.970	0.000	0.589	31
Constant	10.28	0.000		

Table 6: **Estimated relationship between PPP bias and the logarithm of EC income.** The table shows estimation results from regressing PPP bias against the logarithm of EC income. Weights equal to the inverse of the variance of the PPP bias are used. *As we do not have a variance for the United Kingdom, this country is dropped from the estimation and thus we have 31 observations.

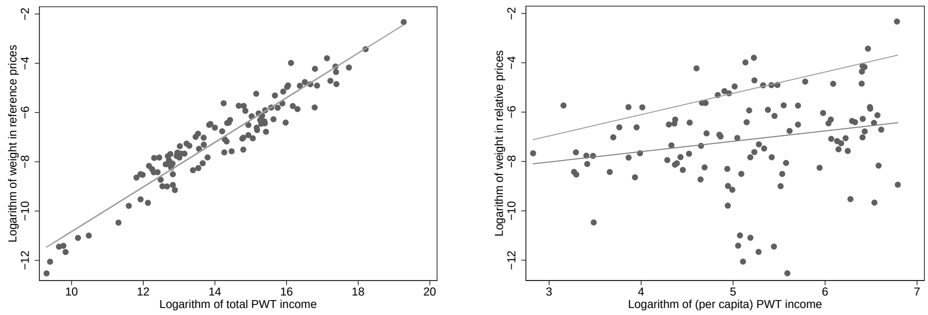

Figure 1: **Weight in the construction of PWT reference prices as a function of PWT income.** The figure displays the logarithm of the difference between the Geary–Khamis reference prices constructed by including all countries, and the reference prices constructed by including all countries but country j, by the logarithm of country j's *total* PWT income (left panel) and by the logarithm of *per capita* PWT income (right panel). The lines display the fitted relationship we obtain when regressing the logarithm of per capita income on the weight in the relative prices. The upper line in the right panel represents the regression giving each country a weight equal to its population size.

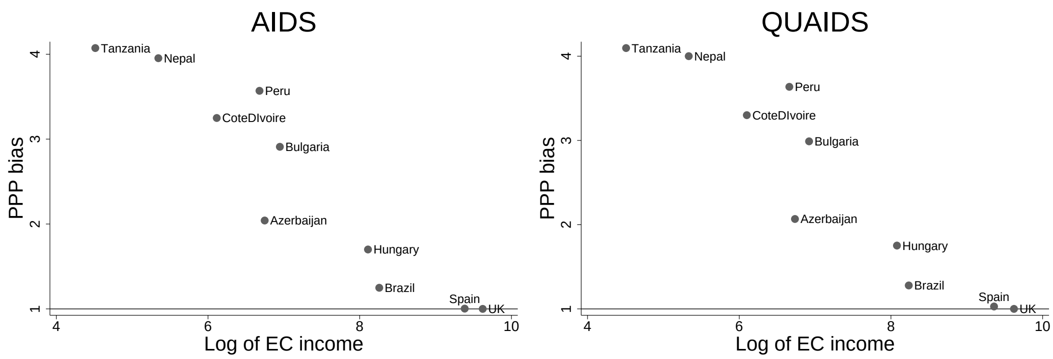

Figure 2: **PPP bias and EC income.** The figure displays the relationship between the estimated PPP bias and EC income for the two different demand systems. The estimates are based on the subsample of households with two children and two adults. The reference line indicates unbiased PWT income relative to the United Kingdom.

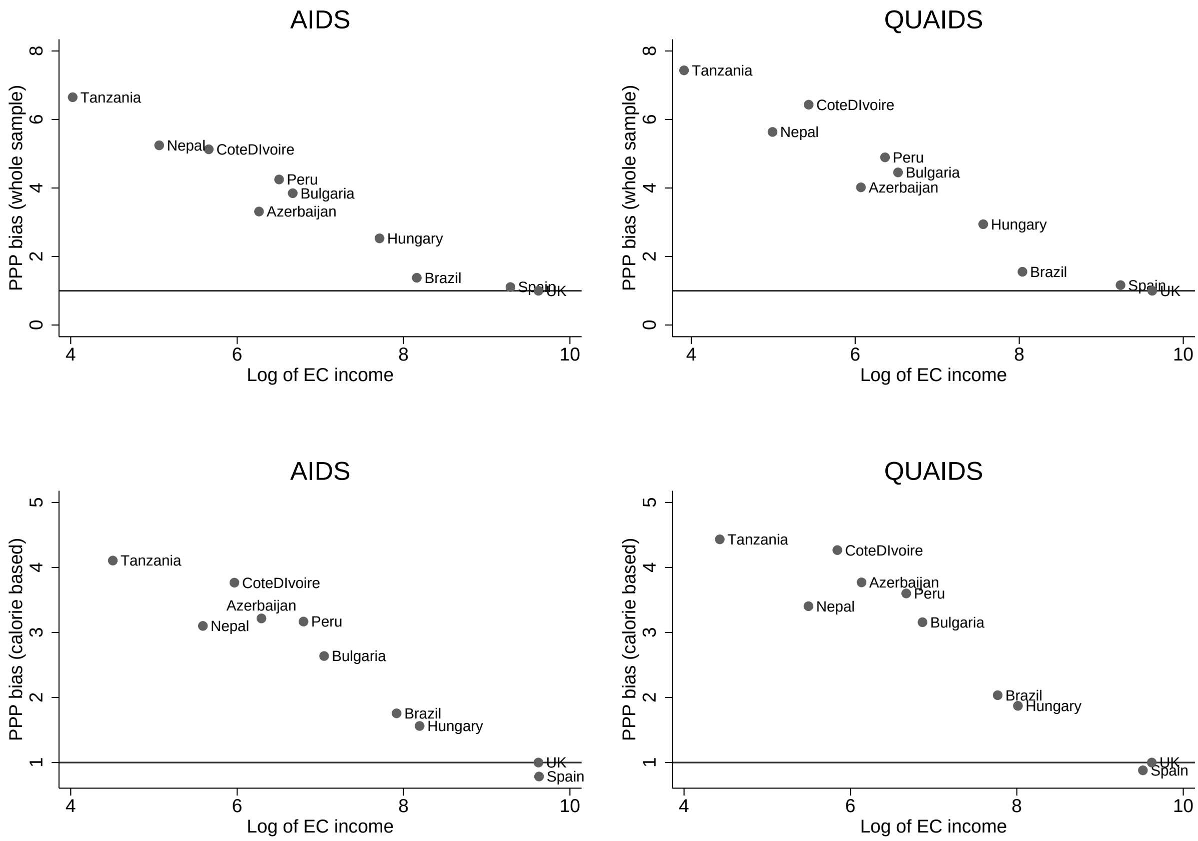

Figure 3: **Robustness analysis.** The figure displays the relationship between the estimated PPP bias and EC income for the two different demand systems. The first row displays the relationship estimated on all households whereas the second row displays the relationship based on the calorie Engel curve. The reference line indicates unbiased PWT income relative to the United Kingdom.

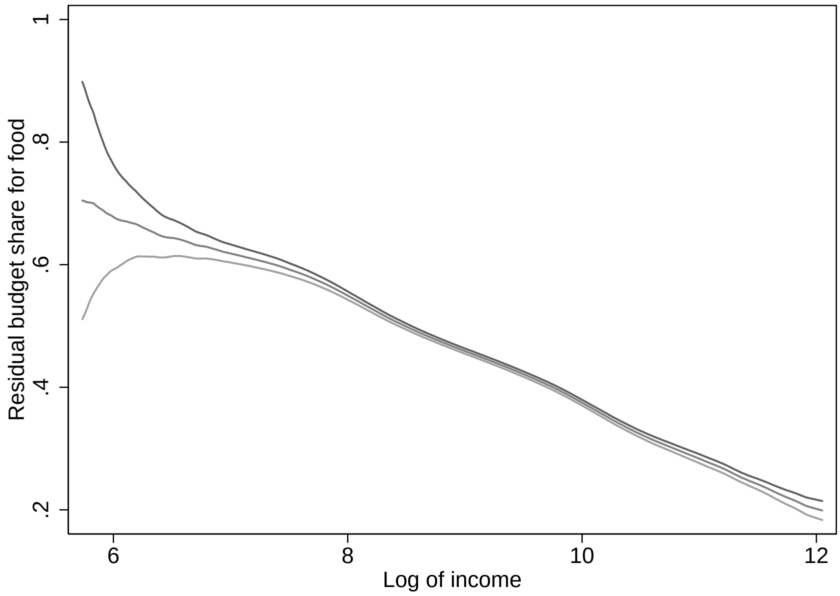

Figure 4: **Kernel regression.** The figure displays the kernel using the Epanechnikov kernel smoother and including households with two children and two adults. The kernel displays the relationship between the budget share for food and the logarithm of household income when the effects of the other explanatory variables are removed by differencing. Tenth-order differencing is conducted based on the optimal differencing weights proposed in Yatchew (2003). The bandwidth is obtained from the formula $bandwidth = 0.15 * (max(logofincome) - min(logofincome))$. The bounds correspond to the 95% confidence intervals. The United Kingdom is used as the base country.

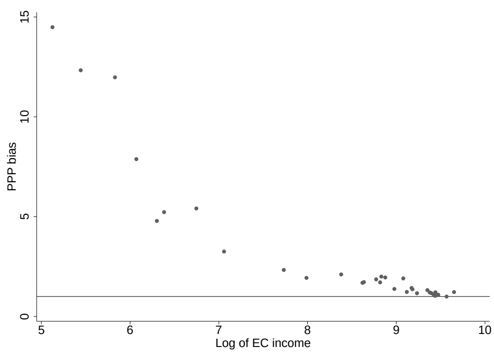

Figure 5: **The relationship between the PPP bias and EC income – extended analysis.** The figure illustrates the relationship between PPP bias and EC income based on the 32 observations in the extended analysis. The reference line indicates the PPP bias level where PWT income is unbiased relative to the United Kingdom.

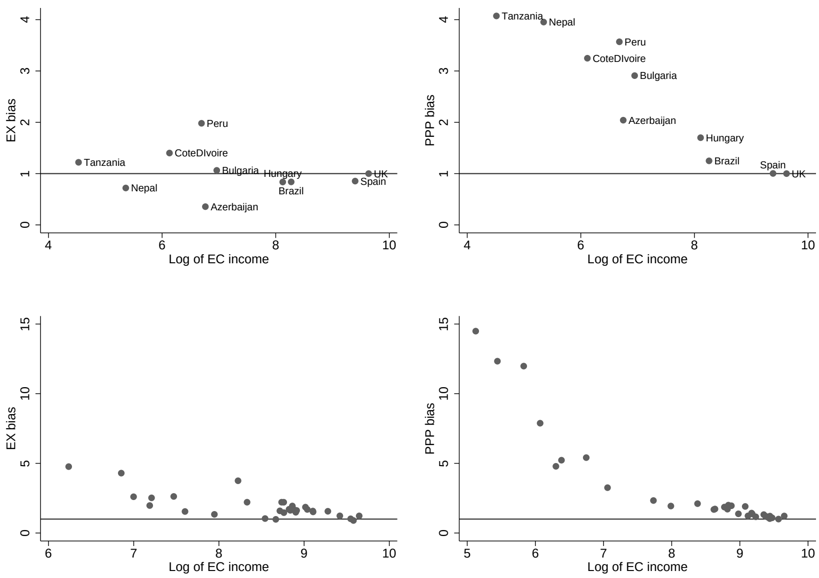

Figure 6: **The relationship between the EX bias, PPP bias and EC income.** The two panels on the left display the PPP bias using the EX method, whereas the two panels on the right display the PPP bias using the PWT price deflator. The two upper figures display the relationship between PPP bias and income in the base model, whereas the lower figures display the same relationship for the extended analysis.

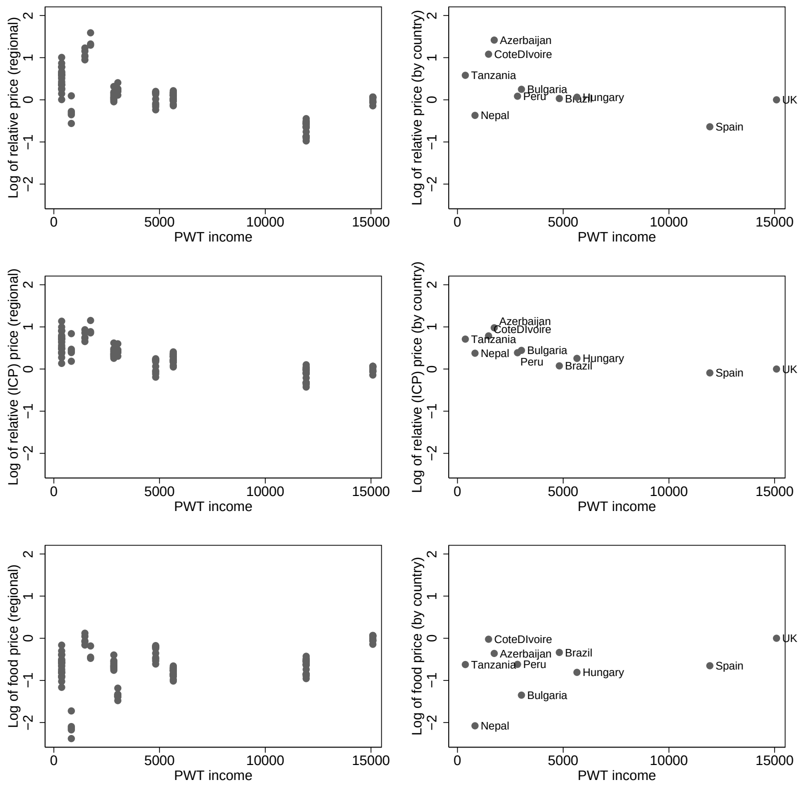

Figure 7: **The relationship between the relative prices and PWT income.** The first row displays the relative prices used in the main analysis, the second row displays the relative prices based on ICP prices, and the third row displays the food prices.

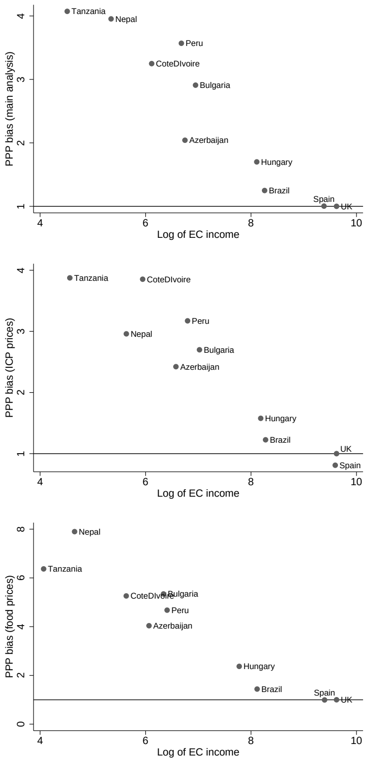

Figure 8: **The relationship between PPP bias and EC income.** The first row shows the relationship between measured PPP bias and EC income in our main analysis, the second row shows the same relationship using ICP prices, and the third row shows this relationship using the food price as a measure of relative price.

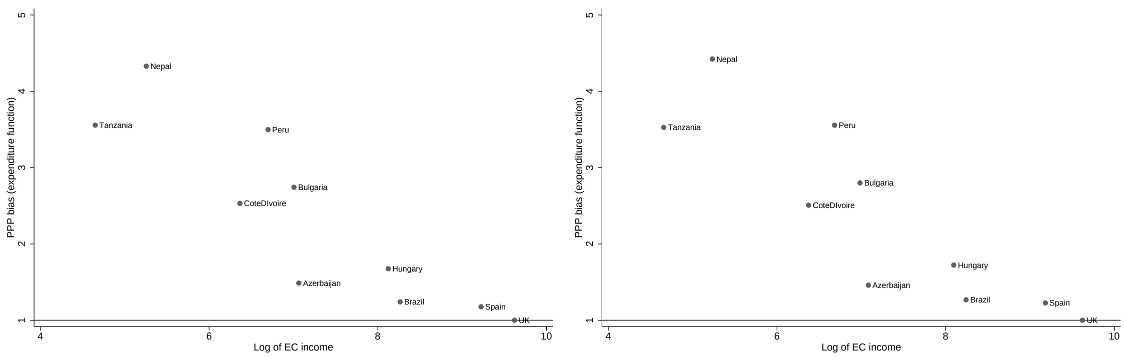

Figure 9: **The relationship between the PPP bias and EC income.** The log of the expenditure function of the AIDS is given by: $ln\_exp_j = log(P_j) + ub_j$ and the log of the expenditure function of the QUAIDS is given by $ln\_exp_j = log(P_j) + ub_j/(1 - ul_b)$, where u is reference utility and b and l are price indexes that are homogenous of degree zero in prices. The PPP bias is given by: exp_j/P'_j. The United Kingdom mean utility level is used as reference utility.