The Price of Development:
the Penn-Balassa-Samuelson effect revisited*

Fadi Hassan

Abstract:

The Penn-Balassa-Samuelson effect is the stylized fact about the positive correlation between cross-country price level and per-capita income. This paper provides evidence that the price-income relation is actually non-linear and turns negative in low income countries. The result is robust along both cross-section and panel dimensions. Additional robustness checks show that biases in PPP estimation and measurement error in low-income countries do not drive the result. The different stage of development between countries can explain this new finding. The paper shows that a model linking the price level to the process of structural transformation captures the non-monotonic pattern of the data. This provides additional understanding of real exchange rate determinants in developing countries.

Keywords: Penn effect; Balassa-Samuelson hypothesis; developing countries; non-parametric estimation; purchasing power parity; real exchange rate; structural transformation.

JEL Classifications: F3, F4, O11.

*Department of Economics, Trinity College Dublin; email: fhassan@tcd.ie; phone: +353 1 896 1667. I have greatly benefited from comments and discussions with Francesco Caselli, Bernardo Guimaraes, Philip Lane, Guy Michaels, Branko Milanovic, Rachel Ngai, Daniel Sturm, Silvana Tenreyro, Adrian Wood, and Alwyn Young. All remaining errors are mine.

1 Introduction

It is widely understood that market exchange rates do not give accurate measures of real income in different economies and that adjustment by purchasing power parity (PPP) factors is necessary for such measures. This understanding is based on an observed empirical regularity that richer countries have a higher price level than poorer countries.[1] The positive correlation between cross-country price level and per-capita income is generally regarded as a stylized fact. This result was documented for twelve developed countries in the seminal paper of Bela Balassa (1964), was confirmed for a large sample of countries as soon as data from the International Comparison Program (ICP) became available and is now renowned as the Penn-Balassa-Samuelson effect (Penn- BS).[2] [3]

The paper makes an important qualification to this general understanding. Using non-parametric estimation, it provides evidence that the price-income relation is non-linear and turns negative in low-income countries both along a cross-section and a panel dimension. Standard regression analysis in subsamples of poor, middle-income, and rich countries is consistent with this finding. The results of the paper are robust to possible sources of bias from PPP estimation and measurement error in low-income countries.

[1]Adjustment by PPPs is necessary as long as price levels vary across countries even if the variation is not systematic with income.

[2]The Penn-BS effect was documented also by Summers and Heston (1991), Barro (1991), and Rogoff (1996). Samuelson (1994) stresses that the proper name for it would be *Ricardo-Viner-Harrod-Balassa-Samuelson-Penn-Bhagwati-et al. effect.*

[3]The Penn-BS effect should not be confused with the Balassa-Samuelson hypothesis The latter provides the mainstream explanation for the former. The Balassa-Samuleson hypothesis argues that richer countries have a higher relative productivity in the tradable sector; under certain assumptions, this leads to a higher relative price of non-tradables, hence to a higher aggregate price level.

The paper argues that the non-monotonicity of the price-income relation is due to the different stages of development that characterize low- and high-income countries. The paper presents a model with three sectors (agriculture, manufacturing, and services) tracing the effects of agricultural productivity, sectoral expenditure and employment shares on the price level of low-income countries. This model captures the non-monotonic pattern of the data, in a way that the standard Balassa-Samuelson hypothesis, focused on productivity differences between tradables and non-tradables, does not. The intuition is that, when a poor country starts to develop, its productivity growth relies mainly in the agricultural sector. Since that, at an early stage of development, agriculture is mainly non-tradable and represents a big share of expenditure, this productivity growth reduces the relative price of agricultural goods, hence the overall price level.

In economics empirical regularities are rare and important. As Solow (1956) and Easterly and Levine (2001) point out, economists build models to match relevant empirical regularities and they use these models to understand economic events and give policy suggestions. The Penn-BS effect is the empirical regularity that the seminal models of Balassa (1964) and Samuelson (1964) try to reproduce. The mechanisms of these models are at the basis of our understanding of long-run real exchange rate movements, are incorporated into many new open-economy macroeconomic models, and have been the initial point of reference for a vast literature on this subject.[4] The paper shows that the empirical regularity that models in the literature are supposed to match, namely the Penn-BS effect, is not actually present in low

[4] The Balassa-Samuelson hypothesis hits more than 7,000 entries on Google Scholar; see Rogoff (1996) and Taylor and Taylor (2004) for extended reviews and Bordo et al. (2014) and Berka et al. (2014) for the most recent applications at the time of writing.

income countries.[5]

The paper makes a significant empirical contribution by uncovering a twist to what has long been accepted as a well-established empirical regularity and offers a novel explanation on real exchange rate determinants in low income countries, based on the process of structural transformation. From a policy point of view, by showing that in poor countries the price-income relation is negative, the paper suggests that there is a "natural" depreciation of the real exchange rate along the development process. This is an important finding that central banks and governments of low-income countries should take into account as they pursue their exchange rate policy. Moreover, the result of the paper suggests that current measures of real exchange rate undervaluation based on the Balassa-Samuelson hypothesis are biased for developing countries; for instance, once we account for the non-monotonic pattern of the price-income relationship, the Chinese Renminbi is 30% less undervalued than standard estimates suggest. The new empirical regularity shown by the paper and its explanation can help us to better understand long-run real exchange rate movements in developing countries and lay the ground for further research on this subject.

The paper relates to the literature on PPPs and the Penn-BS effect as in Kravis, Summers, and Heston (1982), Heston and Summers (1992), and Feenstra et al. (2013). Our contribution is to identify the non-monotonic pattern of the price-income relation as a novel stylized fact. The paper refers

[5]This can explain why there is not much evidence of the Balassa-Samuelson hypothesis in lower income countries (Choudhri and Khan, 2005; Genius and Tzouvelekas, 2008). Notice that they focus on the effect of relative productivity in the tradable sector on the real exchange rate; at the best of our knowledge, the focus on the Penn effect of this paper is a novel contribution.

to the debate on PPPs and real exchange rate determinants in the long run as in Samuelson (1964), Balassa (1964), Bhagwati (1984), Rogoff (1996) and Taylor and Taylor (2004). Within this literature the papers closer in spirit to our are Bergin et al. (2006) and Devereux (1999). The former shows that there is no Penn-BS effect before the '70s; the latter presents a model of endogenous productivity growth in the distribution sector to explain real exchange rate depreciation in east Asian countries. The paper provides a more generalized and systematic evidence of a counter Penn-BS effect and real exchange rate movements in developing countries, explaining the finding with the process of structural transformation. Moreover, Feenstra et al. (2013) argue that the results of Bergin et al. (2006) are driven by interpolation issues of PPPs to past data; this critique does not apply to this paper because our main results are based on a cross-section dimension in benchmark years. Finally, the paper is complementary to the literature on structural transformation and the role of agriculture as a driver of development as in Gollin et al. (2002, 2007) and Ngai and Pissarides (2007). We show the effect that structural transformation out of agriculture has on the real exchange rate in developing countries.

The paper is structured as follows: Section 2 shows that the price-income relation is non-monotonic using both non-parametric and linear estimations. Section 3 establishes that the results are robust to measurement error, bias in the estimation of PPPs, and different databases. Section 4 argues that differences in economic structure can explain the results, derives a model that links the price level to the process of structural transformation, and analyzes the empirical prediction of the model showing that it can capture the non-monotonicity of the data. Section 5 concludes by summarizing the main findings and discussing further research based on these results.

2 The price-income relation

In this section I show that the price-income relation is non-monotonic. I provide evidence along a cross-section and panel dimension through both linear and non-linear estimation. Following the literature on the Penn-BS effect, I measure income per capita in purchasing power parity (PPP) and define the price level as the ratio of the PPP to the exchange rate with the US dollar.[6] Unless alternatively specified, the database of reference is the Penn World Table (PWT) 8.0 version.[7]

2.1 Cross-section dimension

In Figure 1.1, we can see an example of the little attention that the literature has paid to the Penn-BS effect in developing countries. The figure illustrates the positive price-income relation reported in the review of the purchasing power parity puzzle by Rogoff (1996). Since observations with an income per capita lower than Syria are gathered in a cloud of points, it is difficult to properly disentangle the relation between price and income in poor countries.

Therefore, in Figure 1.2, using the same data-set as in Rogoff (1996), I plot the log values of income per capita.[8] I investigate the price-income relation

[6] I use income per capita at constant prices for the panel and time-series analysis and income at current prices for the cross-section analysis.

[7] The results presented in the paper hold also for the World Development Indicators database of the World Bank. I work with the Penn World Tables because traditionally it is the database of reference for this literature.

[8] This is Penn World Table 5.6 (reference year 1985); he considers the year 1990

using a non-parametric estimation technique known as LOWESS (locally weighted scatter smooth), which allows me to impose as little structure as possible on the functional form. This estimation suggests that the Penn-BS effect does not hold in the poorest 25 percent of countries in the sample, where the relation is actually downward sloping. The minimum point of the curve corresponds to an income level of around 1350 PPP $ (1985 prices), which is equivalent to the income of Senegal in the year 1990.

In commenting the result of Figure 1.1, Rogoff (1996) stressed that "*The relation between income and prices is quite striking over the full data set (...); it is far less impressive when one looks either at the rich countries as a group, or at developing countries as group.* In this paper we take Rogoff's point further using a non-parametric estimation that shows that the relation is actually striking when looking at rich countries as a group and negative when looking at poor countries as a group. According to our knowledge, the non-monotonicity of the price-income relation has not been previously documented in the literature.

The LOWESS estimation works as follows: Consider an independent variable x_n and a dependent variable y_n. For each observation y_n the LOWESS estimation technique runs a regression of x_n using few data points around x_n. The regression is weighted so that the central point $(x_n; y_n)$ receives the highest weight and points further away get less weight. The fitted value of this regression evaluated at y_n represents the smoothed value y_n^S which is used to construct the non-parametric curve that links y and x. The procedure is repeated for each observation $(x_n; y_n)$. The number of regressions is equal to the number of observations, and the smoothed curve is the set of all $(x_n; y_n^S)$.

LOWESS estimation requires that the bandwidth of observations included in the regression of each point be chosen. Specifying a large bandwidth provides a smoother estimation, but increases the risk of bias by including observations from other parts of the density. A small bandwidth can better identify genuine features of the underlying density, but increases the variance of the estimation. In the paper I use the default STATA bandwidth of 0.8, which is a conservative choice and provides a lower-bound of the non-monotonic pattern of the data.[9]

Next, I extend the analysis to the PWT 8.0 using only the benchmark countries and the benchmark year.[10] Arguably, this is the best available sample of countries for running this exercise. PWT 8.0 relies on the 2005 ICP round, which provides the most exhaustive dataset for international comparison of real income and prices; moreover, using only the benchmark countries and year minimizes the source of measurement error.[11]

In Figure 2.1 I can confirm the strong positive relation predicted by the Penn-BS effect by running a standard linear estimation of price on income: the OLS coefficient is 0.21 with a t-statistic of 9.23.[12]

[9] The Pseudo-R^2 of the LOWESS estimation is maximized a at a bandwidth of 0.4, which delivers a stronger non-monotonicity at the cost of higher variance. Using a Kernel estimation rather than a LOWESS conveys very similar results to the ones presented in the paper.

[10] I exclude countries with less than one million people in the year 2000 and Zimbabwe and Tajikistan which are clear outliers; including these countries would reinforce the findings. The list of the countries included can be found in the appendix.

[11] All the results presented in the paper hold also using PWT 7.0 or older versions; results are not particularly affected by the upgrade from PTW 7.0 to PWT 8.0; see Feenstra et al. (2013) for a description of the new PWT.

[12] I run an OLS regression, with robust standard errors, of the log of the price level of GDP (variable $pl - gdpe$ from PWT 8.0) and the log of GDP per capita in PPPs at

However, once I allow for non-linearities, the Penn-BS effect breaks down for low income countries. Figure 2.2 shows the results of running a LOWESS estimation between price and income imposing little restriction on the functional form. We can see that the expected upward sloping relation holds only for middle- and high-income countries. The relation is downward sloping for low-income countries; this involves 20 percent of the countries in the sample. The turning point is at 1,448 PPP $ per-capita (2005 prices) equivalent to the income of Senegal in the year 2005. The countries on the downward sloping path are listed in Table 1; we can notice that these are mainly African and Asian (no Latin-American).

Figure 3 reports 95% confidence bands of the LOWESS estimation derived from the standard errors of the smoothed values. The confidence interval confirms the non-monotonic pattern of the data. The Pseudo-R^2 of the non-parametric estimation is 0.6, which is higher than the 0.44 R^2 of the linear model. The F-test comparing the non-parametric model to the linear one rejects the null hypothesis that the non-linear model does not provide a statistically significant better fit.

Standard cross-country OLS regression supports the finding of the non-parametric estimation. In Table 2, I divide the sample by income groups according to standard World Bank classification. The price-income relation is negative, sizable, and significant for low-income countries; it is not statistically different from zero for the middle-income group; and it turns positive

current prices (*cgdpe/pop* from PWT 8.0). I use the expenditure-side of real GDP and price because of comparability with past versions of Penn World Tables and previous studies; the results of the paper are robust to using the measures of real GDP and price from the output-side newly introduced by PWT 8.0.

and significant for high-income countries. The results of the OLS regressions are consistent with the non-monotonicity of the price-income relation stressed by the non-parametric estimation.[13]

Moreover, Table 3 shows that a quadratic specification of the price-income relations confirms the non-monotonic pattern. Both $Income$ and $Income^2$ are statistically significant. The coefficient associated to the linear term is negative and the quadratic one is positive, indicating a convex relation. The marginal effect of income on price turns positive around 1,800 PPP $ per-capita (2005 prices), which is equivalent to the income of Laos in the year 2005. The turning point from the quadratic specification is at a higher level of income than from the previous non-parametric estimation.

Given the functional form $Price_i = \alpha + \beta\, Income_i + \gamma\, Income_i^2 + \epsilon_i$, Lind and Mehlum (2011) show that in order to test for the presence of a U-relation, it is necessary to formulate the following joint null hypothesis:

$$H_0 : \beta + 2\,\gamma\, Income_{min} \geq 0 \; and/or \; \beta + 2\,\gamma\, Income_{max} \leq 0 \qquad (1)$$

against the alternative:

$$H_1 : \beta + 2\,\gamma\, Income_{min} < 0 \; and \; \beta + 2\,\gamma\, Income_{max} > 0 \qquad (2)$$

Lind and Mehlum (2011) build a test for the joint hypotheses using Sasabuchi's (1980) likelihood ratio approach. Table 4 shows that the the marginal effect of income on price is negative and statistically significant at $Income_{min}$ and

[13] The observation per-income group are 36, 58, and 32 respectively. The World Bank threshold is 875 US$ (2005) for low-income countries and 10,276 US$ (2005) for middle-income countries.

positive ans statistically significant at $Income_{max}$. The bottom line of the table shows that the SLM test rejects H_0 in favor of the alternative and thus indicates that the result is consistent with the presence of a U-relation between price and income.

2.2 Panel dimension

In this section, I analyze the price-income relation in a panel dimension. The ICP collects data prices only in benchmark years. Then, the PWTs used to estimate prices for other years by rescaling according to the inflation rate differential with the US. The new version of the Penn World Tables makes use of historical ICP benchmarks to extrapolate the time series of prices and real incomes. This new data set relies on a better methodology. However, many countries, especially developing ones like China or India, did not participate to all the benchmark collections; this makes the computation of prices and real incomes in non-benchmark years more uncertain. Nevertheless, PWTs are regularly used in empirical analyses with panels; moreover, panel regressions of price on income are commonly used to build measures of real exchange rate over/undervaluation. Thus, it is relevant to assess if the non-monotonicity of the price-income relation holds also along a panel dimension.

If I extend the analysis to a panel of countries between 1950-2009, standard linear estimation of price on income confirms the positive relation predicted by the Penn-BS effect: the OLS coefficient is 0.15 with a t-statistic of 32.7 (Figure 4.1).[14] However, non-parametric estimation shows that the price-

[14]This is for a sample of 126 countries from 1950 to 2009 using PWT 8.0. Countries with less than one million people in the year 2000. I run an OLS regression of the log

income relation is also non-monotonic along a panel dimension. The Penn-BS effect holds for middle- and high-income countries, but in low-income countries the relation is negative (Figure 4.2).

Figure 4.3 reports the fitted value of the LOWESS estimation. The turning point is at 1421 PPP $ per-capita (2005 prices), which corresponds to the income of Senegal in the year 2000. The downward sloping arm of the curve includes 27% of the total observations, and 45% of the countries in the sample. The countries on the downward sloping arm and their frequencies are reported in Table 5. We can see that some of the countries are persistently on the downward-sloping arm (i.e. Ethiopia and Tanzania); others moved along the curve (i.e. China and Vietnam).

Standard panel-data analysis, Table 6, confirms the result of the non-parametric estimation. I take 5-years averages of price and income between 1950-2009. I show that for developing countries the relation between price and income is negative and significant with and without country fixed-effects. I do this by running a regression for the full sample, and then for developing countries only.[15] This result comes despite a broad definition of developing countries and a linear restriction on the price-income relation.

of the price level of GDP (variable pl_gdpe) and the log of GDP per capita in PPPs at constant chained prices ($rgdpe/pop$).

[15] I define developing countries those below the World Bank's threshold of high-income countries; a stricter definition of developing countries reinforces my result. Notice that in the full sample with country fixed effects the coefficient is not significantly different from zero.

3 Robustness checks

The data used to estimate the price-income relationship are PPPs, exchange rates , and GDP per-capita.[16] Most of the robustness analysis focuses on PPPs by looking at measurement error in prices and at bias in the construction of PPPs, which are arguably the main source of concern. Moreover, given that in developing countries official exchange rates can be different from black market rates, I control for this possible source of bias. Finally, I show that results are robust to different versions of the Penn World Tables.[17]

3.1 Classical measurement error

Chen et al. (2007) analyze the bias of the OLS estimation of price on income when there is measurement error in prices. In this case the independent variable becomes correlated with the error term, so that the standard assumptions for a consistent and unbiased least square estimator break down.[18] Chen et al. (2007) conclude that the OLS estimate will be biased downwards and can become negative if the variance of the measurement error is

[16] I remind the reader that in the Penn World Tables $p = \frac{PPP}{XRAT}$ and $y = \frac{GDP}{PPP}$

[17] I do not focus the robustness discussion on estimates of GDP per-capita. Gollin et al. (2014) analyze the definitions and measurement approaches used in the construction of national accounts data in poor countries. They conclude that these aggregate data are robust to problems associated with informality or household production and that there is no reason to believe that they are intrinsically flawed.

[18] The econometric specification of the price-income relation is such that $p_i^* = \alpha + \beta y_i^* + \epsilon_i$, where variables are expressed in logs and p_i^* is the true price level without measurement error and $y_i^* = Y_i - p_i^*$ is the true real income per-capita. Consider the case where the measured price level p_i contains an error such that $p_i = p_i^* + \eta_i$, where η_i has mean zero and is normally distributed; then the regressor and the error term become correlated.

sufficiently high. In fact, they show that:[19]

$$\text{plim } \hat{\beta} = \frac{\beta^{true} - \frac{\sigma_\eta^2}{\sigma_{y^*}^2}}{1 + \frac{\sigma_\eta^2}{\sigma_{y^*}^2}} \tag{3}$$

where σ_η^2 is the variance of measurement error and $\sigma_{y^*}^2$ is the variance of the true real income per-capita. From this expression we can see that as the variance of the measurement error σ_η^2 increases, the estimated $\hat{\beta}$ can turn negative.

In the group of low-income countries the OLS estimate of price on income is -0.21 (Table 2). What is the level of measurement error's variance needed to drive this result? Assuming that measurement error is correlated to the level of income but not to the level of price, we can rewrite equation (3) as:[20]

$$\text{plim } \hat{\beta} = \frac{\beta^{true} - \frac{\sigma_\eta^2}{\sigma_{y^*}^2}}{1 + \frac{\sigma_\eta^2}{\sigma_{y^*}^2}} = \frac{\beta - \frac{\sigma_\eta^2}{\sigma_Y^2 + \sigma_p^2 + \sigma_\eta^2 - 2\sigma_{Yp}}}{1 + \frac{\sigma_\eta^2}{\sigma_Y^2 + \sigma_p^2 + \sigma_\eta^2 - 2\sigma_{Yp}}} \tag{4}$$

In the sub-sample of countries where the price-income relation is negative, we have $\sigma_Y^2 = 1.8$, $\sigma_p^2 = 0.18$, $\sigma_{Yp} = 0.66$ (remember that all the variables are expressed in logs).

The variance of measurement error that would lead to the negative estimation of -0.21 depends on the value of β^{true}. Let's suppose that β^{true} is equal to the OLS estimation over the full sample (0.21). In this case, in order to

[19] Assuming that the measurement error is uncorrelated with the true dependent and independent variables as well as with the equation error, equation (3) follows.

[20] From the specification of Chen et al. (2007), we have that $y_i^* = Y_i - p_i + \eta_i$; keeping the same independence assumptions of their paper, equation (4) follows.

get $\hat{\beta} = -0.21$, we would need $\sigma_\eta^2 = 0.74$: the measurement error on prices should have a variance 4 times higher than the variance of observed prices over the full sample. If we rather assume that in the group of low income countries β^{true} is zero, we would need $\sigma_\eta^2 = 0.42$: hence in this case the variance of the measurement error on prices in this sub-sample of countries should be more than double than the variance of the observed prices.

Therefore, even if measurement error could potentially drive the results of the paper, an improbable high variance of the measurement error itself is required to obtain the negative price-income relation presented in the paper.

3.2 Purchasing power parities bias

The process of computing PPPs is subject to intrinsic fragilities, making the comparison of real income and prices across countries a difficult exercise. The PWTs rely on data collected by the International Comparison Program (ICP). In each country the ICP calculates prices for about 155 goods, called basic headings, by collecting prices for 1500-2000 items following a standardized product description (SPD).[21] A basic heading is the most disaggregated level at which expenditure data are available from national accounts statistics. The ICP collects quotes for different items within each basic heading and then computes a unique price through different procedures.[22] Once the prices of all 155 goods are obtained, the PWTs compute

[21] A SPD lists the characteristics relevant to a particular cluster of products and they are elaborated at a regional level with the collaboration of national statistical offices. An example of SPD is: " Men's shirt, well known brands, 100% cotton, light material, classic styling, uniform colour, short sleeves, classic collar, buttons fastner" (ICP, 2007). The ICP regions are Africa, Asia-Pacific, CIS, South America, OECD-Eurostat, Western Asia.

[22] For instance, for the basic heading *rice*, the ICP collects quotes for six different kinds of rice, including long-grained, short-grained, and brown rice. The country-product-dummy regression is the method mostly used to obtain a unique price of the basic heading rice.

a PPP index for each country following the Geary-Khamis (GK) method of aggregation, which compares domestic prices with world prices. In the GK method the world price of a good is defined as a weighted average of its price in all countries and the weights are given by a country's share in the global consumption of that good.

All this process generates various potential sources of bias in the estimation of PPPs. The main ones are: the GK method of aggregation of basic headings into the PPP index; quality matching; and items' representativity (Deaton and Heston, 2010; ICP, 2007). The direction of the PPP bias can have a key influence on our results. Let's suppose that the true price-income relationship is flat. Figure 5 shows that if in low-income countries PPPs tend to be overestimated a negative price-income relationship would arise because of that bias; however, if PPPs are underestimated, a Penn-BS effect would emerge.[23]

The literature has well established that PPPs in low-income countries are underestimated (Nuxoll, 1994; Neary, 2004, Hill, 2004; Deaton and Heston, 2010; Almas, 2012). This implies that the negative price-income relationship in poor countries shown in the paper is a lower bound of the true one.

The GK method of aggregation understates PPPs in poor countries. In fact, countries with a larger physical volume of consumption get a greater weight in the construction of world prices. This implies that the vector of international prices used as a reference is closer to the price of rich rather

See Rao (2004) for a detailed explanation of the items' methods of aggregation

[23]The underlying assumption of Figure 5 is that PPPs bias affects mostly poorer countries.

than poor countries.[24] This generates a Gershenkron effect for low income countries according to which PPP is lower the more the price of a country differs from the price of reference (Gershenkron, 1947; Nuxoll, 1994). This effect stems from the substitution bias that characterizes indexes with a single reference price vector as in the GK method. These type of indexes do not account for utility maximizing agents switching towards cheaper goods as relative prices change (Hill, 2000).[25]

The method of aggregation is not the only source of bias of PPPs. Quality matching is also a problem because the estimation of PPPs makes use of a set of homogeneous goods. As Deaton and Heston (2010) stress, one of the most criticized issues of ICP rounds is that lower quality goods and services in poor countries are often matched to higher quality items in rich countries. Quality mismatch leads to an underestimation of the price level in poor countries; hence also this source of bias reinforces the results of the paper.

Finally the representativity of the items whose prices are collected is also a potential source of bias. This relates both to the aggregation of items into a basic heading and to the urban bias in collecting prices. If an item within the basic heading is representative in some countries but not in others, PPPs

[24] Nuxoll (1994) shows that international prices are closest to that of a moderately prosperous country like Hungary.

[25] Neary (2004) shows that the GK method of aggregation is exact if preferences are Leontief; in this case goods are perfect complements and the substitution bias does not arise. Different methods of aggregation like the Elteto-Koves-Szulc index (EKS) used by the World Bank mitigates the Gershenkron effect for poorer countries. The PPP-EKS index of a country takes a geometric mean over all the possible Fisher indexes of all the countries with both the country in question and a reference country (for a discussion comparing the two methods see Deaton and Heston, 2010). Using the PPP-EKS index reinforces the non-monotonicity shown in the paper; results available upon request.

may be estimated incorrectly.[26] This is a common problem for all ICP rounds.[27] Nevertheless, Diewert (2008) argues that if non-representative prices are well-distributed across all countries in a region, they may not cause serious distortions. Moreover, Deaton (2010) computes a Tornqvist index to measure how much different goods moves the overall PPP-index in Africa and Asia.[28] He concludes that there is no evidence to support the idea that prices in Africa or in the Asia-Pacific region are systematically overstated by the representativity issue.

Feenstra et al. (2013) show that in China the price level has been overstated because of a urban bias in the data collection. In order to account for this bias the PWT introduces a uniform reduction of 20% to the ICP prices. This adjustment is consistent with their estimates of China's real GDP. Our results account for this downward revision. However, there is no clear evidence of price overestimation for other countries due to the urban bias. Actually Atkin and Donaldson (2012) show that the price of detailed products in Ethiopia and Nigeria are on average 5-12% higher in rural areas. Therefore, urban bias should not be driving the results of the paper.

To summarize, the method of aggregation and quality matching tend to bias downwards the estimation of PPPs in low-income countries compared the "true" values. Moreover, there is no evidence that products representativity systematically biases PPPs upwards or that the urban bias affects the coun-

[26] See for example the wheat versus teff example in Deaton and Heston (2010).

[27] The Latin American region tried to overcome this issue in the 2005 round by using an extended CPD method, adding a representativity dummy. The OECD/Eurostat and CIS regions used an EKS method based on Javon indexes of representative products between countries; see ICP (2007) for a brief description of this method.

[28] He estimates a pairwise Tornqvist index for the ring African countries vs. the UK and at regional level for Africa and Asia-Pacific vs. OECD/Eursotat.

tries on the downward sloping path of the price-income relation. Therefore, the non-monotonicity shown in Section 2 is actually a lower-bound.

3.3 Previous versions of the Penn World Tables and black market exchange rates

The analysis of the paper makes use of the Penn World Table 8.0 database. This relies on the 2005 ICP round, which arguably provides the best available data for international comparisons of real income. The PPPs of many developing countries were revised upwards after this round, and these countries have a lower real income than what was previously thought (Deaton, 2010). Although higher PPPs in poor countries work in favor of my findings, the last ICP round does not drive the results of the paper and they hold also for previous versions of the PWTs.

In Figure 6, I run a series of cross-section LOWESS estimations of the price-income relation for benchmark years and benchmark countries of subsequent versions of the PWT.[29] The non-monotonicity of the price-income relation is confirmed also for these older versions of the PWT.[30] Moreover, it is interesting to observe that the relative income of the turning point of the relation decreases over time, so we observe an increasing Penn-Balassa-Samuelson effect as stressed by Bergin et al. (2006).

Another potential issue to account for is that the PWTs use official exchange rates to compute the price level, but in developing countries the official

[29] I use PWT 5.6 for 1985, PWT 6.1 for 1996, and PWT 7 for 2005

[30] The non-monotonicity holds also for the panel dimension; results available upon request

rates can greatly differ from the one actually used in daily transactions, above all in the early years of our sample. Nevertheless, this issue does not undermine the finding of the paper. As Reinhart and Rogoff (2004) argue, multiple exchange rate arrangements decreased greatly over time and apply mainly until the 1980s, while the non-monotonicity of the price-income relation shown in the paper takes the year 2005 as a benchmark. However, I have run a non-parametric estimation of price on income using black market exchange rates for the year 1996 and the non-monotonicity of the relation is confirmed also in this case.[31]

This section has shown that the results of the paper are robust to classical measurement error, bias in PPPs estimation, that they hold for different versions of the PWTs and are not affected by using black market exchange rates. All this provides evidence that the non-monotonicity of the price-income relation is not a spurious result, but a hitherto-undocumented economic fact.

4 Theoretical explanation

4.1 Beyond the Balassa-Samuelson hypothesis

The most accepted explanation of the Penn-BS effect is the Balassa-Samuelson (BS) hypothesis. This explanation focuses on productivity differentials between the tradable and the non-tradable sector. Assuming free labor mobility across sectors and that the law of one price holds for tradables, the

[31] I choose the year 1996 because this is the oldest benchmark year for which both black market rates and raw PPPs are available. Results available upon request. Data on black market rates are taken from Reinhart and Rogoff (2004). Prices are computed dividing PPPs from PWT 6.1 by the black market exchange rates.

BS hypothesis shows that countries with higher relative productivity in the tradable sector have a higher price level. Since richer countries tend to have higher relative productivity in the tradable sector, the price level should then raise with per-capita income. [32]

In order to capture the non-monotonicity of the price-income relation, this paper argues that we need a modified BS framework that accounts for the relevance of the agricultural sector in poor countries and for the fact that low-income and high-income countries have very different economic structures and are at different stages of development. In Table 7, I consider the benchmark countries of PWT for the year 2005. I rank countries by their level of income and divide the sample by terciles. Then, following the tradition of the development macroeconomics literature, I focus on a sectoral division of the economy between agriculture, manufacturing, and services.

We can see that the countries in the bottom group of income have a remarkably different structure in terms of valued added, expenditure, and employment shares. The most significant differences refer to the agricultural sector: the group of countries where the price-income relation is negative have a 10 times higher valued-added share in agriculture, a 5 times higher expenditure share and a 9 times higher employment share than the countries in the top group of income. This clearly reflects the early stage of development that characterizes these countries.

If structural change is an important determinant of the non-monotonicity of

[32] Devereux (1999) shows that a counter Penn-BS effect can arise if there is a higher productivity growth in the non-tradable sector for instance due to improvements in the distribution of the service sector. A higher productivity in the non-tradable sector and a reclassification of the non-tradable sectors are key in this paper.

the price-income relation, we should observe some non-monotonicity when the price level is regressed against key indicators of structural change. Figure 7 confirms that this is the case by showing a non-monotonic pattern of the price level respect to both employment and expenditure shares in agriculture.

The differences in value added, expenditure and employment shares are associated to a different structure of relative prices. Using disaggregated data kindly provided by the International Comparison Program at the World Bank, I can compute sectoral PPPs and price levels.[33] Perhaps contrary to conventional wisdom, the relative price of agriculture in terms of both services and manufacturing turns to be higher in low-income countries than in rich-countries. [34] Moreover, the average price level of services and manufacturing increases by income group, but the price level of agriculture decreases between the bottom and the intermediate group. Non parametric estimations of sectoral prices on income confirm this pattern: Figure 8 shows that the price dynamics of the agricultural sector accounts for most of the non-monotonicity of the overall price-income relation.

The explanation for the non-monotonic price-income relation that this paper proposes is therefore the following: When a poor country starts to develop, its productivity growth relies mainly in the agricultural sector. This allows for a reduction of the relative price of agricultural goods. Since in a country

[33]The price level of sector i is given by $p_i = PPP_i/XRAT$ with $p_i^{US} = 1$. In order to preserve aggregation at the GDP level, I use the Geary-Khamis method to compute sectoral PPPs. See the appendix A.5 for a detailed description of sectoral classification of goods; as suggested by Herrendorf and Valentinyi (2011), I map the agricultural sector with the food sector.

[34]Caselli (2005) hints at this possibility in a footnote. Lagakos and Waugh (2012) have a similar finding.

at an early stage of development, agriculture represents a big share of both expenditure and value added production, there is an overall reduction of the price level. After a certain stage of development the share of the agricultural sector in the economy decreases. Hence the previous effect fades out and productivity gains from the manufacturing sector becomes a more important source of growth, so that we are back to the standard Balassa-Samuelson mechanism.

The two key elements of this explanation are that productivity growth in the agricultural sector is higher than in other sectors and that agricultural goods are not tradable. Duarte and Restuccia (2010) show for a panel of 29 countries between 1956-2004 that productivity growth was 4% in agriculture, 3% in manufacturing and 1.3% in services. As for the non-tradability of agricultural goods, this is a plausible assumption for low-income countries; Gollin et al. (2007) argue that *" it is reasonable to view most [poor] economies as closed from the perspective of trade in food"*. They show that in the year 2000 about 70% of arable land in 159 developing countries was devoted to staple food crops. With the exception of few developing countries, almost all of the resulting production was for domestic consumption. Moreover, food imports and food aid are not a major source of food for poor countries: imports of food supply around 5% of total calories consumed.

4.2 Structural change and the price level

This section develops a model that links the price level of a country to its process of structural transformation. It derives a consumption-based price index from the utility function, within a modified version of the Balassa-Samuelson framework. Then, taking as reference Ngai and Pissarides (2007), it expresses the consumption shares of that index as a function of the employment shares. In this way the price level can reflect a country's stage of development. In the next section, I then test if the price implied by this model can generate a non-monotonic price-income relation.

Production functions are given by:

$$F_i(k_i, l_i) = A_i k_i^{\alpha} n_i^{1-\alpha}; \quad i = a, m, s \tag{5}$$

Factors' market clearing satisfy:

$$\sum_{i=1}^{m} l_i = 1; \quad \sum_{i=1}^{m} k_i = k; \tag{6}$$

Moreover, we have that $F_i = c_i$ for $i = a, s$. We also assume that manufacturing produces both a final consumption good and the economy's capital stock so that $\dot{k} = F^m - c_m - (\delta + n)k$.

The underlying assumptions of the model, as in the Balassa-Samuelson framework, are that manufacturing is the only tradable and that trade is balanced period by period. These imply that the effect of trade is to equalize the price of manufacturing across countries and that there is financial autarky, which is a reasonable assumption for low-income countries. The purpose of these assumptions is to have a model as close as possible to the standard Balassa-Samuelson framework.

The utility function is assumed to have constant elasticities across goods so that:

$$U(c_a, c_m, c_s) = \left[\gamma_a^{\frac{1}{\theta}} c_a^{\frac{\theta-1}{\theta}} + \gamma_m^{\frac{1}{\theta}} c_m^{\frac{\theta-1}{\theta}} + \gamma_s^{\frac{1}{\theta}} c_s^{\frac{\theta-1}{\theta}} \right]^{\frac{\theta}{\theta-1}} \tag{7}$$

The consumption-based price index P is defined as the minimum expenditure:

$$z = P_a c_a + P_m c_m + P_s c_s \tag{8}$$

such that $c = U(c_a, c_m, c_s) = 1$ given P_i.

So defined, the consumption-based price index measures the least expenditure that buys a unit of the utility-based consumption index. Under standard assumptions, the consumption-based price index can be written as: [35]

$$\log P = \gamma_a \log p_a + \gamma_s \log p_s; \tag{9}$$

As in Ngai and Pissarides (2007), we can link the expenditure shares to the employment shares, so that the price index can be finally expressed as: [36]

$$\log P^{BS+} = (\gamma_a + \gamma_s) \left[\log A_m - \left(\frac{l_a}{l_a + l_s} \log A_a + \frac{l_s}{l_a + l_s} \log A_s \right) \right] \tag{10}$$

where l_i is the employment share of sector i, A_i is TFP in sector i. I label it Balassa-Samuelson+ price index.

[35]See the appendix A.2 for a complete derivation.
[36]See the appendix A.3 for a complete derivation.

4.3 Alternative predictions of the Price-Income relation

In this section I compute the price level implied by the standard Balassa-Samuelson hypothesis and by the "Balassa-Samuleson+" hypothesis. I then use these price levels to estimate the price-income relation non-parametrically and compare the fitted values with the actual pattern of the data.

Under the Balassa-Samuelson hypothesis, the price level of country z is:

$$\log P^{BS} = \gamma_{NT}(\log A_T - \log A_{NT}) \tag{11}$$

where γ_{NT} is the expenditure share of non-tradables.

Notice that (10) and (11) are very similar. The differences are that in the Balassa-Samuelson+ there is a better focus on the agricultural sector and the sectoral TFPs of agriculture and services are weighted by the relative employment shares, so that the price index reflects the stage of structural transformation. If we shut down the focus on the agricultural sector by setting γ_a and l_a equal to zero, as if it were absorbed by the manufacturing sector, we are back to the standard Balassa-Samuelson hypothesis.

In order to compute these price levels, I obtain sectoral estimates of TFP across countries following the method of Herrendorf and Valentinyi (2011).[37] Employment shares are taken by the WDI database and by national sources. The saving rate σ is set equal to the share of investment in GDP. The consumption share in agriculture and service are given by the expenditure

[37]They elaborate a development accounting framework to compute sectoral productivities using the Penn World Tables; see the appendix for a detailed description

shares from the ICP database.[38]

Figure 9.1 shows the fitted values of the non-parametric estimation of the price-income relation, where prices are given by equation (11): I am able to confirm the strictly positive relation predicted by the Balassa-Samuelson hypothesis.

However, Figure 9.2 shows that the price implied by the "BS+" hypothesis allows for more flexibility in the price-income relation and can generate a negative pattern at low levels of development. Therefore, by taking into account that countries are at a different stage of their process of structural transformation, I am able to match better the actual pattern of the data reported in Figure 9.3.

Table 8 analyzes the quantitative fit: under the BS+ hypothesis 26 percent of countries in the sample are on the downward sloping path of the price-income relation; in the standard BS hypothesis this is 0% and in the actual data it is 20% of the sample. The variance of prices generated by the BS+ hypothesis is two and half times higher than in the data (1.02 vs 0.41). Finally, the turning point of the BS+ model is around 3,000 PPP$, but in the data it is around 1,440 PPP$.

The quantitative result of the "Balassa-Samuleson+" hypothesis clearly outperforms that of the Balassa-Samuelson hypothesis. The model derived in this paper is relatively simple and a richer approach that accounts for other

[38] I am able to compute the price levels for 60 countries out of 127 because of the lack of sectoral employment data in many poor countries and lack of investment data in middle-income and former URSS countries; following Caselli (2005) I exclude countries with data on investment starting only after the '70s.

factors like the tradability of agriculture in rich countries or the reduction of trade costs as a country develops might deliver a better quantitative fit. However the results presented are encouraging and lay the ground for further theoretical and empirical research on the relation between structural change and the price level.

5 Conclusions

In this paper I show that the relation between price and income is non-monotonic. To my knowledge, this is an original finding and it is a hitherto undocumented empirical regularity. This result contradicts the conventional wisdom of a positive price-income relation, which draws upon a linear estimation. If I apply a non-parametric estimation, the price-income relation turns out to be significantly negative in poor countries. This finding is robust along both cross-section and panel dimensions. The new evidence presented in this paper raises general questions about the relation between the process of economic development and the price level, as well as about the long-run determinants of real exchange rate in poor countries.

The paper shows that a model linking the price level to the process of structural transformation that characterizes developing countries can generate a non-monotonic pattern of the price-income relation. This result suggests that structural change and, more in general, inter-sectoral dynamics can be important determinants of real exchange rates movements. Nevertheless, a richer theoretical approach could improve the quantitative fit.

For instance, the model does not account for the role of trade costs. Trade costs are much higher than is generally recognized, even for traded goods: Anderson and Van Wincoop (2004) estimate that, for developed countries, trade costs average 170% of production costs, of which roughly half is international trade costs and half internal trade costs. For developing countries, they claim that this ratio is often higher, and many studies do indeed show strikingly high transport costs for individual developing countries or groups thereof (Limao and Venables, 2001).

Trade costs and the ratio of trade costs to production costs may well vary systematically with the level of development. For example as a low-income country starts developing, its infrastructure improves reducing both internal and external trade costs as well as the ratio of trade costs to production. This might turn to be a key element in explaining the initial negative pattern of the price-income relation and deserves further investigation. This is consistent with Du et al. (2013) who show that transport infrastructure is an important determinant of exchange rate especially in developing countries.

The tradability of agriculture in more developed countries is another feature that a richer model should account for. In the current model, agriculture is completely non-tradable and this could partly explain the high variance of prices and the turning point's high level of income that the model predicts.

Finally, a possible empirical extension of the paper could focus on regional variation within countries like India or China, where there are regions at very different stages of development. This kind of regional variation would be ideal to verify if the process of structural transformation is at the basis of the non-monotonic price-income relation.

This paper lays the ground for further theoretical and empirical research on the relation between economic development and the price level. The results presented, although surprising, should not be disturbing. It is probable that Samuelson himself would not have been startled. In his 1994 article for the thirty-year anniversary of the Balassa-Samuelson model, he wrote that " *The Penn-Balassa-Samuelson effect is an important phenomenon of actual history but not an inevitable fact of life. It can quantitatively vary and,* ***in different times and places, trace to quite different processes***".

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A Appendix

A.1 Countries in the cross-section analysis of section

Albania	Cote d'Ivoire	Japan	Netherlands	Sweden
Angola	Croatia	Jordan	New Zealand	Switzerland
Argentina	Czech Republic	Kazakhstan	Niger	Syria
Armenia	Denmark	Kenya	Nigeria	Taiwan
Australia	Ecuador	Korea	Norway	Tanzania
Austria	Egypt	Kuwait	Oman	Thailand
Azerbaijan	Estonia	Kyrgyzstan	Pakistan	Togo
Bangladesh	Ethiopia	Laos	Paraguay	Tunisia
Belarus	Finland	Latvia	Peru	Turkey
Belgium	France	Lebanon	Philippines	Uganda
Benin	Gabon	Lesotho	Poland	Ukraine
Bolivia	Gambia, The	Liberia	Portugal	United Kingdom
Bosnia and Herz.	Georgia	Lithuania	Romania	United States
Botswana	Germany	Macedonia	Russia	Uruguay
Brazil	Ghana	Madagascar	Rwanda	Venezuela
Bulgaria	Greece	Malawi	Saudi Arabia	Vietnam
Burkina Faso	Guinea	Malaysia	Senegal	Yemen
Cambodia	Guinea-Bissau	Mali	Serbia	Zambia
Cameroon	Hong Kong	Mauritania	Sierra Leone	
Canada	Hungary	Mauritius	Singapore	
Central Afr. Rep.	India	Mexico	Slovak Rep.	
Chad	Indonesia	Moldova	Slovenia	
Chile	Iran	Mongolia	South Africa	
China	Iraq	Morocco	Spain	
Colombia	Ireland	Mozambique	Sri Lanka	
Congo, Dem. Rep.	Israel	Namibia	Sudan	
Congo, Rep. of	Italy	Nepal	Swaziland	

A.2 The Consumption-Based Price Index

The consumption-based price index P is defined as the minimum expenditure:

$$z = P_a c_a + P_m c_m + P_s c_s \tag{12}$$

such that $c = \phi(c_a, c_m, c_s) = 1$ given P_i.

So defined, the consumption-based price index measures the least expenditure that buys a unit of the consumption index on which period utility depends.

From consumer's utility maximization we know that:

$$\frac{MU_i}{MU_j} = \frac{P_i}{P_j} \tag{13}$$

so that:

$$\left(\frac{\gamma_a}{\gamma_m}\right)^{\frac{1}{\theta}} \left(\frac{c_m}{c_a}\right)^{\frac{1}{\theta}} = \frac{P_a}{P_m}; \quad c_a = \frac{\gamma_a}{\gamma_m} c_m \left(\frac{P_a}{P_m}\right)^{-\theta} \tag{14}$$

and

$$\left(\frac{\gamma_s}{\gamma_m}\right)^{\frac{1}{\theta}} \left(\frac{c_m}{c_s}\right)^{\frac{1}{\theta}} = \frac{P_s}{P_m}; \quad c_s = \frac{\gamma_s}{\gamma_m} c_m \left(\frac{P_s}{P_m}\right)^{-\theta} \tag{15}$$

Substituting c_a and c_s from (14) and (15) into (12) we have:

$$z = \frac{P_a^{1-\theta}}{P_m^{-\theta}} \frac{\gamma_a}{\gamma_m} c_m + P_m c_m + \frac{P_s^{1-\theta}}{P_m^{-\theta}} \frac{\gamma_s}{\gamma_m} c_m \tag{16}$$

so that rearranging:

$$c_m = \frac{\gamma_m P_m^{-\theta} z}{\gamma_a P_a^{1-\theta} + \gamma_m P_m^{1-\theta} + \gamma_s P_s^{1-\theta}} \tag{17}$$

and consequently:

$$c_a = \frac{\gamma_a P_a^{-\theta} z}{\gamma_a P_a^{1-\theta} + \gamma_m P_m^{1-\theta} + \gamma_s P_s^{1-\theta}} \tag{18}$$

$$c_s = \frac{\gamma_s P_s^{-\theta} z}{\gamma_a P_a^{1-\theta} + \gamma_m P_m^{1-\theta} + \gamma_s P_s^{1-\theta}} \tag{19}$$

Equations (17), (18), and (19) are the demands that maximize c given spending z. The highest value of the utility function c given z, thus is found by substituting these demands into (7):

$$\left[\gamma_a^{\frac{1}{\theta}} \left(\frac{\gamma_a P_a^{-\theta} z}{x}\right)^{\frac{\theta-1}{\theta}} + \gamma_m^{\frac{1}{\theta}} \left(\frac{\gamma_m P_m^{-\theta} z}{x}\right)^{\frac{\theta-1}{\theta}} + \gamma_s^{\frac{1}{\theta}} \left(\frac{\gamma_s P_s^{-\theta} z}{x}\right)^{\frac{\theta-1}{\theta}}\right]^{\frac{\theta}{\theta-1}} \tag{20}$$

where $x = \gamma_a P_a^{1-\theta} + \gamma_m P_m^{1-\theta} + \gamma_s P_s^{1-\theta}$.

Since P is defined as the minimum expenditure z such that $c = 1$ we have:

$$\left[\gamma_a^{\frac{1}{\theta}} \left(\frac{\gamma_a P_a^{-\theta} P}{x}\right)^{\frac{\theta-1}{\theta}} + \gamma_m^{\frac{1}{\theta}} \left(\frac{\gamma_m P_m^{-\theta} P}{x}\right)^{\frac{\theta-1}{\theta}} + \gamma_s^{\frac{1}{\theta}} \left(\frac{\gamma_s P_s^{-\theta} P}{x}\right)^{\frac{\theta-1}{\theta}}\right]^{\frac{\theta}{\theta-1}} = 1 \tag{21}$$

from which the solution for P is:

$$P = \left(\gamma_a P_a^{1-\theta} + \gamma_m P_m^{1-\theta} + \gamma_s P_s^{1-\theta}\right)^{\frac{1}{1-\theta}} \tag{22}$$

This is the consumption-based price index consistent with the CES utility function specified in equation (3). When $\theta = 1$ the utility function becomes Cobb-Douglas; in this case the price index becomes:

$$\log P = \frac{\log(\gamma_a P_a^{1-\theta} + \gamma_m P_m^{1-\theta} + \gamma_s P_s^{1-\theta})}{1-\theta} \tag{23}$$

Applying L'Hopital's rule we have:

$$\lim_{\theta \to 1} \frac{\log(\gamma_a P_a^{1-\theta} + \gamma_m P_m^{1-\theta} + \gamma_s P_s^{1-\theta})}{1-\theta} = \frac{f(\theta)}{g(\theta)} = \lim_{\theta \to 1} \frac{f'(\theta)}{g'(\theta)} = \gamma_a \log P_a + \gamma_m \log P_m + \gamma_s \log P_s \tag{24}$$

so that for the Cobb-Douglas case, the consumption-based price index is given by:

$$\log P = \gamma_a \log P_a + \gamma_m \log P_m + \gamma_s \log P_s \tag{25}$$

Accounting for the cross-country equalization of the price of manufacturing through trade and normalizing it to one, the consumption-based price index can be written as:

$$\log P = \gamma_a \log p_a + \gamma_s \log p_s \tag{26}$$

A.3 Relative prices, consumption shares and employment

From the supply-side, static efficiency condition requires equal marginal rate of technical substitution across sectors, so that $k_i = k$; while free movement of capital and labor leads to equal remuneration of the factors of production. Therefore, firms' profit maximization implies:

$$\frac{P_a}{P_m} = \frac{A_m}{A_a} \tag{27}$$

$$\frac{P_s}{P_m} = \frac{A_s}{A_a} \tag{28}$$

From consumer's optimality conditions (

$$\frac{P_a c_a}{P_m c_m} = \frac{\gamma_a}{\gamma_m}\left(\frac{P_a}{P_m}\right)^{1-\theta} \equiv x_a \tag{29}$$

$$\frac{P_s c_s}{P_m c_m} = \frac{\gamma_s}{\gamma_m}\left(\frac{P_s}{P_m}\right)^{1-\theta} \equiv x_s \tag{30}$$

We then define $X = x_a + x_s + x_m$, where clearly $x_m = 1$. We also define:

$$c \equiv \sum_{i=1}^{m} P_i c_i; \qquad y \equiv \sum_{i=1}^{m} P_i F^i \tag{31}$$

Using equations (29) and (30) and the efficiency conditions, we can rewrite equations (31) as:

$$c = P_m c_m X; \qquad y = P_m A_m k^{\alpha} \tag{32}$$

Notice that the technology parameter for output is TFP in manufacturing not an average of all sectors.

As in Ngai and Pissarides (2007) we can link relative expenditure with the employment shares. If we substitute we substitute $F^i = c_i$ for $i = a, s$ in (29) and (30), using the market clearing conditions in (6), we can show that it results:

$$l_a = \frac{c}{y}\frac{x_a}{X} \tag{33}$$

$$l_s = \frac{c}{y}\frac{x_s}{X} \tag{34}$$

The employment share in the manufacturing sector is derived by firstly observing that $l_m = 1 - l_a - l_s$, so that we have:

$$l_m = \frac{c}{y}\frac{x_m}{X} + \left(1 - \frac{c}{y}\right) \tag{35}$$

Let's consider the case where $\theta = 1$ and manufacturing is the numeraire. In this case the price index is given by $\log P = \gamma_a \log p_a + \gamma_s \log p_s$. By using firm's optimality conditions (27) and (28) as well as (33) and (34) We can write the price level as:

$$\log P = (\gamma_a + \gamma_s)\left[\log A_m - \left(\frac{l_a}{l_a + l_s}\log A_a + \frac{l_s}{l_a + l_s}\log A_s\right)\right] \tag{36}$$

A.4 Sectoral TFPs Methodology

In order to compute sectoral TFPs, I use the methodology of Herrendorf and Valentinyi (2011) who elaborate a sectoral development accounting framework that allows to compute sectoral TFPs using PWT. The key assumptions of their methodology are: competitive markets; factor's mobility across sectors; Cobb-Douglas production function with factor shares common to all countries.

The production function for sector i in country z is given by:

$$y_i^z = A_i^z (k_i^z)^{\theta_i} (l_i^z)^{\phi_i} (h_i^z)^{1-\theta_i-\phi_i} \tag{37}$$

where k is capital, l is land, and h is human capital.

Under the assumption stated above, Herrendorf and Valentinyi (2011) show that the sectoral factors of production are:

$$k_i^z = \frac{\theta_i p_i^z y_i^z}{\sum_j \theta_j p_j^z y_j^z} \sum_i k_i^z \tag{38}$$

$$l_i^z = \frac{\phi_i p_i^z y_i^z}{\sum_j \phi_j p_j^z y_j^z} \sum_i l_i^z \tag{39}$$

$$h_i^z = \frac{(1-\theta_i-\phi_i) p_i^z y_i^z}{\sum_j (1-\theta_j-\phi_j) p_j^z y_j^z} \sum_i h_i^z \tag{40}$$

In order to compute sectoral TFPs, I take the sectoral factor shares from Herrendorf and Valentinyi (2011), who calculate them from the US input-output tables. Then, following their methodology, I compute the capital stock in the economy k^z with the perpetual inventory method as in Caselli (2005). Land l^z is arable land for agriculture and urban land for manufacturing and services. I take data on arable land from FAOSTAT and following World Bank (2006) estimates, I set urban land equal to 24% of physical capital. Finally, I compute human capital h^z as in Caselli (2005) and it is an increasing function of average years of schooling per worker.

A.5 ICP 2005, classification of goods

Category	Basic Heading	*BS-SC framework:* **Sector allocation**	*BS-framework:* **Tradability**
Food	Rice	A	T
	Other cereals and flour	A	T
	Bread	A	T
	Other bakery products	A	T
	Pasta products	A	T
	Beef and veal	A	T
	Pork	A	T
	Lamb, mutton and goat	A	T
	Poultry	A	T
	Other meats and preparations	A	T
	Fresh or frozen fish and seafood	A	T
	Preserved fish and seafood	A	T
	Fresh milk	A	T
	Preserved milk and milk products	A	T
	Cheese	A	T
	Eggs and egg-based products	A	T
	Butter and margarine	A	T
	Other edible oils and fats	A	T
	Fresh or chilled fruit	A	T
	Frozen, preserved or processed fruits	A	T
	Fresh or chilled vegetables	A	T
	Fresh or chilled potatoes	A	T
	Frozen or preserved vegetables	A	T
	Sugar	A	T
	Jams, marmalades and honey	A	T
	Confectionery, chocolate and ice cream	A	T
	Food products n.e.c.	A	T

Category	Basic Heading	*BS-SC framework:* **Sector allocation**	*BS-framework:* **Tradability**
Beverages and tobacco	Coffee, tea and cocoa	M	T
	Mineral waters,soft drinks,fruit and veg juices	M	T
	Spirits	M	T
	Wine	M	T
	Beer	M	T
	Tobacco	M	T
Clothing and footwear	Clothing materials and accessories	M	T
	Garments	M	T
	Cleaning and repair of clothing	S	NT
	Footwear	M	T
	Repair and hire of footwear	S	NT
Housing, water, electricity and gas	Actual and imputed rentals for housing	S	NT
	Maintenance and repair of the dwelling	S	NT
	Water supply and miscellaneous services relating to the dwelling	S	NT
	Miscellaneous services relating to the dwelling	S	NT
	Electricity	M	T
	Gas	M	T
	Other fuels	M	T
Furniture, household equipment and maintenance	Furniture and furnishings	M	T
	Carpets and other floor coverings	M	T
	Repair of furniture, furnishings and floor coverings	S	NT
	Household textiles	M	T
	Major household appliances whether electric or not	M	T
	Small electric household appliances	M	T

Category	Basic Heading	*BS-SC framework:* **Sector allocation**	*BS-framework:* **Tradability**
Furniture, household equipment and maintenance	Repair of household appliances	S	NT
	Glassware, tableware and household utensils	M	T
	Major tools and equipment	M	T
	Small tools and miscellaneous accessories	M	T
	Non-durable household goods	M	T
	Domestic services	S	NT
	Household services	S	NT
Health	Pharmaceutical products	M	T
	Other medical products	M	T
	Therapeutical appliances and equipment	M	T
	Medical Services	S	NT
	Dental services	S	NT
	Paramedical services	S	NT
	Hospital services	S	NT
Transport	Motor cars	M	T
	Motor cycles	M	T
	Bicycles	M	T
	Fuels and lubricants for personal transport equipment	M	T
	Maintenance and repair of personal transport equipment	S	NT
	Other services in respect of personal transport equipment	S	NT
	Passenger transport by railway	S	NT
	Passenger transport by road	S	NT
	Passenger transport by air	S	NT

Category	Basic Heading	*BS-SC framework:* **Sector allocation**	*BS-framework:* **Tradability**
Transport	Passenger transport by sea and inland waterway	S	NT
	Combined passenger transport	S	NT
	Other purchased transport services	S	NT
Communication	Postal services	S	NT
	Telephone and telefax equipment	M	T
	Telephone and telefax services	S	NT
Recreation and culture	Audio-visual, photographic and information processing equipment	M	T
	Recording media	M	T
	Repair of audio-visual, photographic and information processing equipment	S	NT
	Major durables for outdoor and indoor recreation	M	T
	Other recreational items and equipment	M	T
	Gardens and pets	S	NT
	Veterinary and other services for pets	S	NT
	Recreational and sporting services	S	NT
	Cultural services	S	NT
	Games of chance	S	NT
	Newspapers, books and stationery	S	NT
	Package holidays	S	NT
Education	Education	S	NT
Restaurant and hotels	Catering services	S	NT
	Accommodation services	S	NT
Miscellaneous goods and services	Hairdressing salons and personal grooming establishments	S	NT
	Appliances, articles and products for personal care	S	NT

Category	Basic Heading	*BS-SC framework:* **Sector allocation**	*BS-framework:* **Tradability**
Miscellaneous goods and services	Prostitution	S	NT
	Jewellery, clocks and watches	M	T
	Other personal effects	M	T
	Social protection	S	NT
	Insurance	S	NT
	FISIM	S	NT
	Other financial services n.e.c	S	NT
	Other services n.e.c.	S	NT
Government expenditure	Government compensation of employees	S	NT
	Government intermediate consumption	M	T
	Government gross operating surplus	S	NT
	Government net taxes on production	S	NT
	Government receipts from sales	S	NT
Capital formation	Metal products and equipment	M	T
	Transport equipment	M	T
	Residential buildings	M	T
	Non-residential buildings	M	T
	Civil engineering works	M	T
	Other products	M	T
Inventories	Changes in inventories and acquisitions	M	T

A=agriculture; M=manufacturing; S=services; T=tradable; NT=non-tradable.

The sectoral allocation and the tradability allocation apply respectively to the estimation of the Balassa-Samuelson-Structural-Change and the Balassa-Samuelson framework in section 4.

Tables

Table 1: Countries on the downward sloping arm of the LOWESS estimation, cross-section dimension

Bangladesh	Liberia
Benin	Madagascar
Burkina Faso	Malawi
Central African Republic	Mali
Congo, Dem Rep.	Mozambique
Ethiopia	Nepal
Gambia	Niger
Ethiopia	Rwanda
Gambia	Sierra Leone
Guinea	Tanzania
Guinea-Bissau	Togo
Kenya	Uganda
Lesotho	Zambia

Table 2: Cross-country OLS regression by income groups, year 2005

Dependent var: ln *price*	ln income
Low income	-0.21**
	(-3.85)
Middle income	0.06
	(0.65)
High income	0.51***
	(2.29)
Full sample	0.21***
	(9.23)

*** Significant at the 1% level; ** significant at the 5% level; robust t-statistics in parenthesis.

Table 3: Cross-country OLS regression: linear and quadratic specifications, year 2005

Dependent var: ln *price*	(1)	(2)
ln *income*	0.21***	-1.61***
	(9.23)	(-7.09)
ln $income^2$		0.11***
		(7.80)
N. Obs.	126	126
R^2	0.44	0.64

*** Significant at the 1% level; robust t-statistics in parenthesis.

Table 4: Tests for a U-shape

Dependent var: ln *price*	
Slope at $Income_{min}$	-0.44***
	(-5.68)
Slope at $Income_{max}$	0.72***
	(9.74)
SLM test for U-shape	5.68
p-value	0.00

*** Significant at the 1% level; robust t-statistics in parenthesis.

Table 5: Countries on the downward sloping arm of the LOWESS estimation, panel dimension

Country	Frequency	Country	Frequency	Country	Frequency
Bangladesh	38	Guinea	24	Nigeria	14
Benin	53	Guinea-Bissau	52	Pakistan	20
Bolivia	7	India	45	Paraguay	5
Bosnia Herzegovina	4	Indonesia	15	Philippines	3
Botswana	16	Iraq	1	Romania	2
Brazil	2	Kenya	24	Rwanda	41
Burkina Faso	53	Korea	14	Senegal	4
Cambodia	35	Laos	24	Sierra Leone	48
Cameroon	15	Lesotho	51	Sudan	33
Central African Rep.	52	Liberia	33	Syria	16
Chad	44	Madagascar	52	Taiwan	2
China	30	Malawi	58	Tanzania	50
Congo, Dem. Rep.	62	Mali	48	Thailand	17
Congo, Republic of	20	Mauritania	26	Togo	52
Cote d'Ivoire	2	Mongolia	13	Tunisia	1
Egypt	34	Morocco	11	Uganda	46
Ethiopia	62	Mozambique	52	Vietnam	11
Gambia	52	Nepal	52	Yemen	15
Ghana	13	Niger	52	Zambia	20

Table 6: Panel evidence on price level and real income, 1950-2009 (5-years average)

Dependent var: ln *price*	**Full Sample**		**Developing Countries**	
	(1)	(2)	(1)	(2)
ln income	0.08***	0.002	-0.11***	-0.18***
	(2.38)	(0.04)	(-2.51)	(-2.79)
Country, fe	NO	YES	NO	YES
Time dummies	YES	YES	YES	YES
No. of countries	126	126	9.4	9.4
Avg obs per country	9.7	9.7	94	94

*** Significant at the 1% level; robust t- and z-statistics in parenthesis.

Table 7: Price-income relation and the stage of development

price-income relation		1st Tercile negative	2nd Tercile flat	3rd Tercile positive
Value-added share of GDP				
	Agriculture	30.46	11.09	2.84
	Manufacturing	26.42	37.00	31.95
	Services	43.12	51.92	65.21
Employment share				
	Agriculture	60.61	28.02	6.65
	Manufacturing	10.50	22.10	26.01
	Services	28.33	49.13	66.97
Expenditure share				
	Agriculture	35.08	20.45	8.47
	Manufacturing	41.71	43.86	41.42
	Services	20.28	25.15	29.91
Price level				
	Agriculture	0.67	0.63	1.06
	Manufacturing	0.56	0.63	1.03
	Services	0.19	0.27	0.77

Table 8: Data and models

	Data	BS+ Model	BS Model
Countries on the downward sloping path	20%	26%	0%
Price, Std. Deviation	0.41	1.02	0.02
Turning point	1,464 PPP$	3,070 PPP	-

Figures

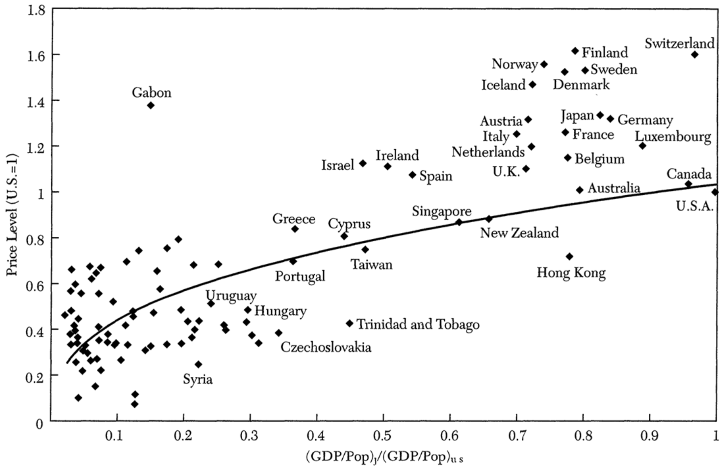

Figure 1.1: Price Level and Income - Rogoff (1996)

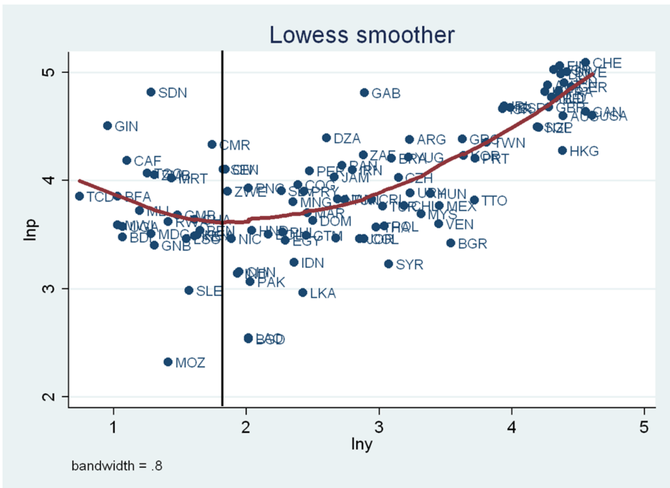

Figure 1.2: Price Level and Income - Rogoff (1996); log-income & non-param. estimation

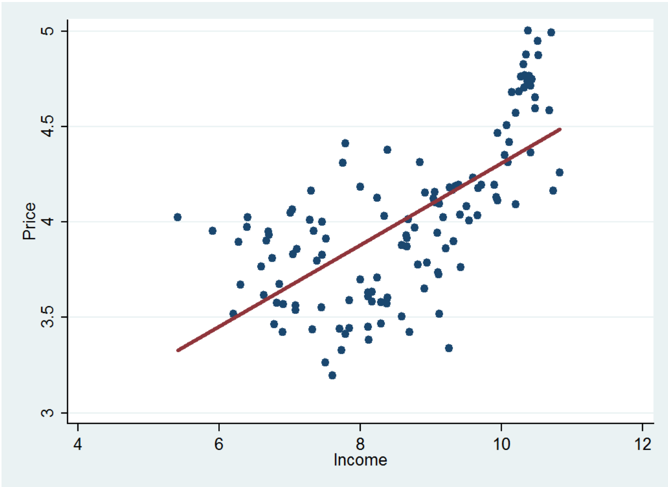

Figure 2.1: Price level and Income PWT 8.0, benchmark countries, 2005: Linear Estimation

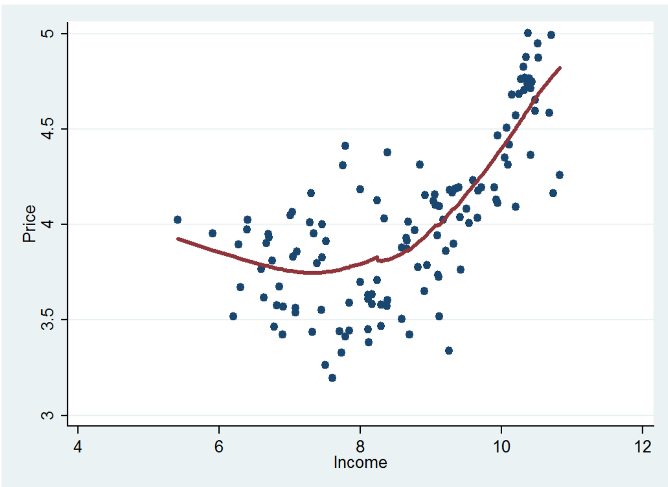

Figure 2.2: Price level and Income PWT 8.0, benchmark countries, 2005: Non-Parametric Estimation

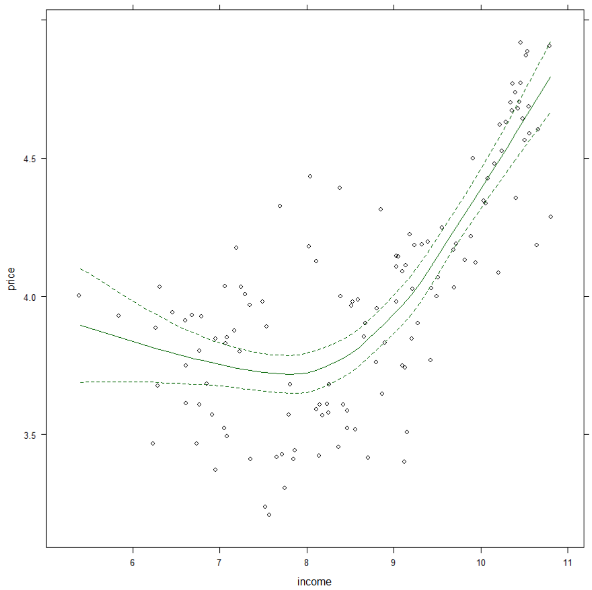

Figure 3: Price and Income PWT 8.0, benchmark countries, 2005: Non-Parametric Estimation, 95% confidence bands

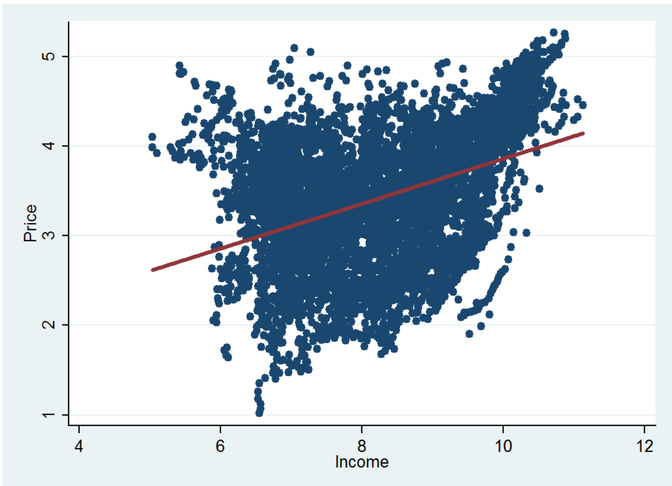

Figure 4.1: Prices and Income 1950-2011: OLS Estimation

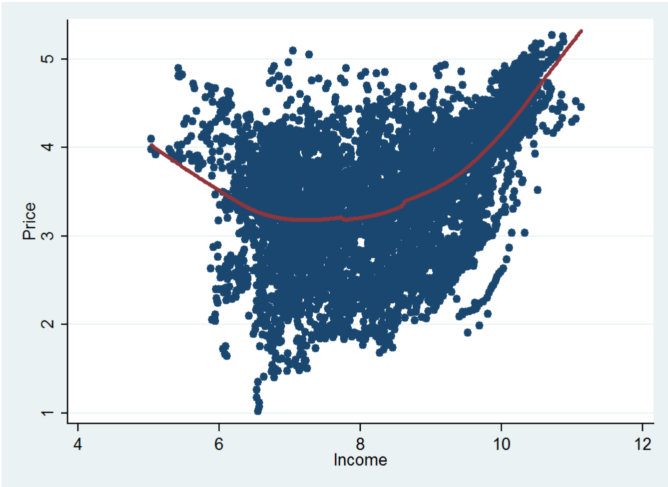

Figure 4.2: Prices and Income 1950-2011: Non-Parametric Estimation

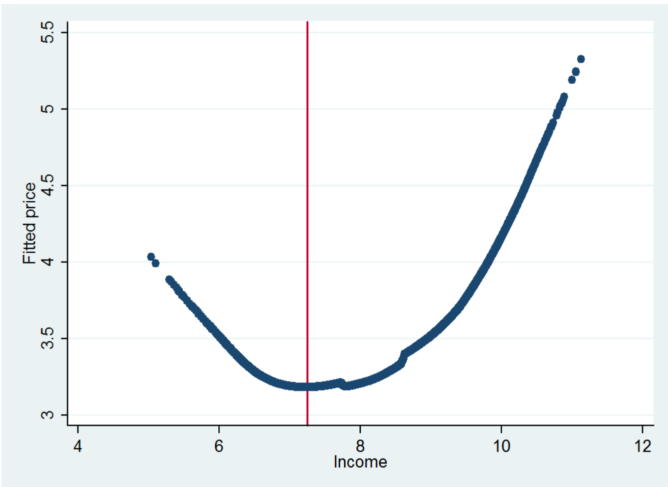

Figure 4.3: Prices and Income 1950-2011: Non-Parametric Estimation, fitted values

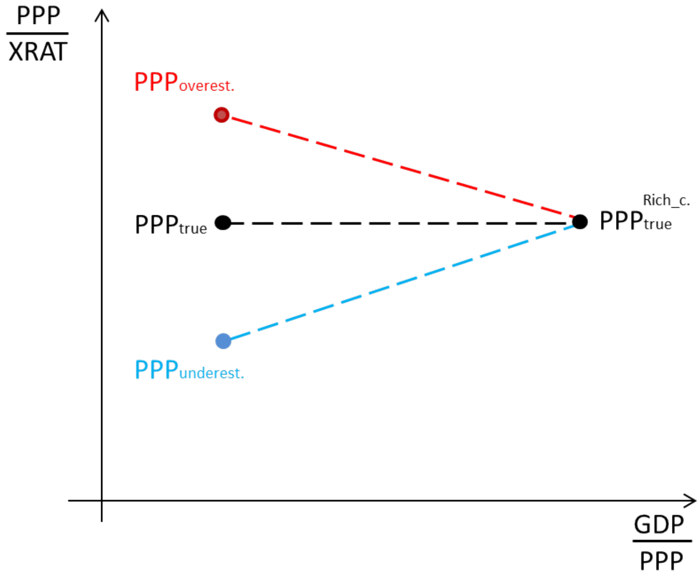

Figure 5: The effect of PPPs bias

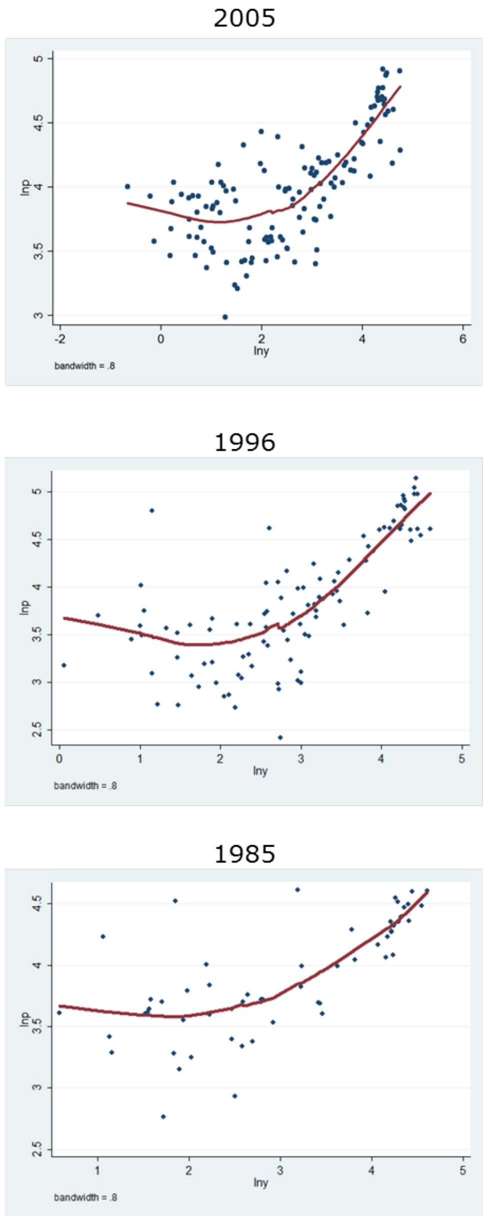

Figure 6: Price and income: benchmark years and countries

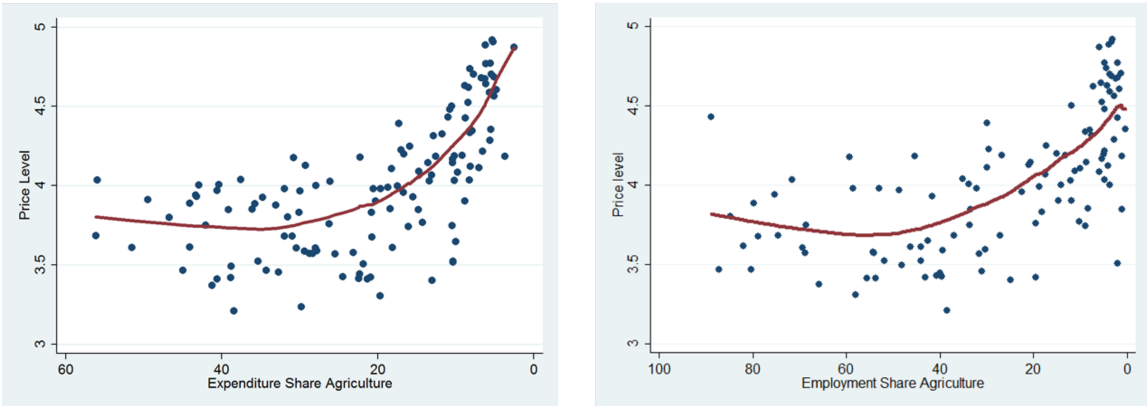

Figure 7: Price Level, Expenditure and Employment Share of Agriculture (reversed scale): Non-Parametric Estimation

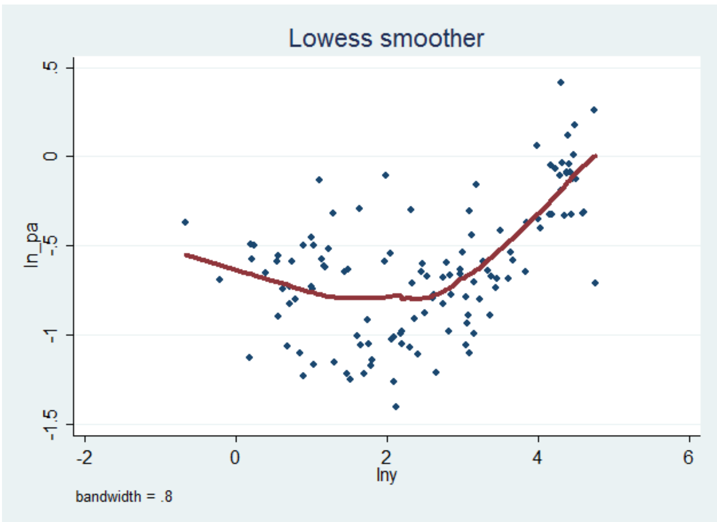

Figure 8.1: Price of Agriculture and Income: Non-Parametric Estimation

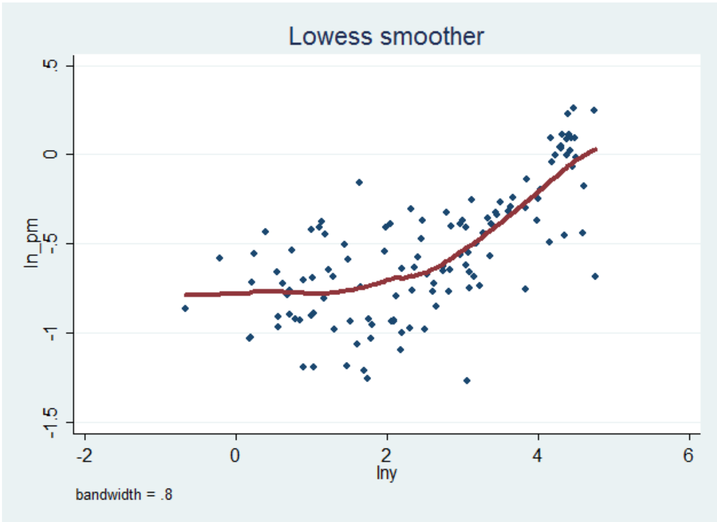

Figure 8.2: Price of Manufacturing and Income: Non-Parametric Estimation

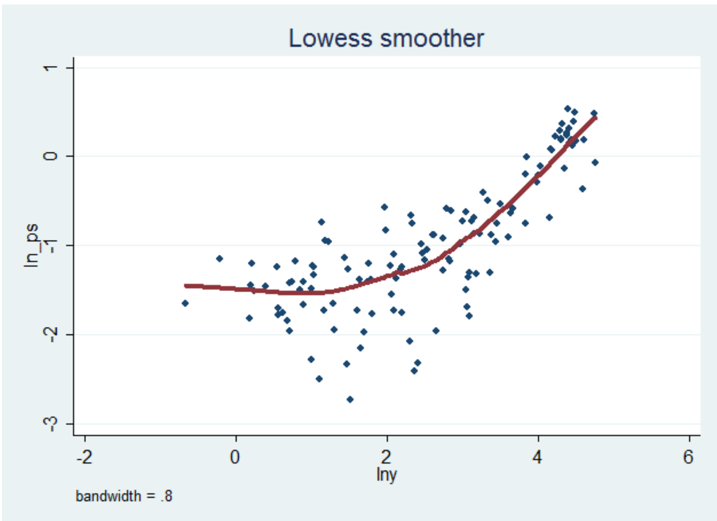

Figure 8.3: Price of Services and Income: Non-Parametric Estimation

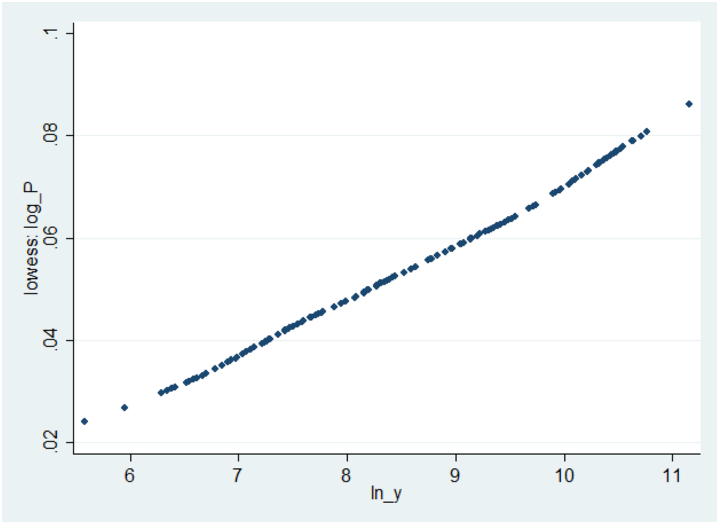

Figure 9.1: The price level in the Balassa-Samuelson hypothesis: non-parametric estimation of the price-income relation, fitted values

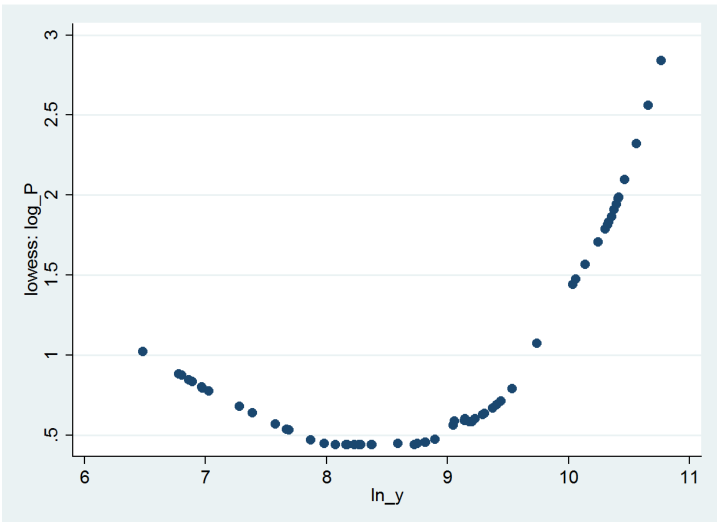

Figure 9.2: The price level in the Balassa-Samuelson+ hypothesis: non-parametric estimation of the price-income relation, fitted values

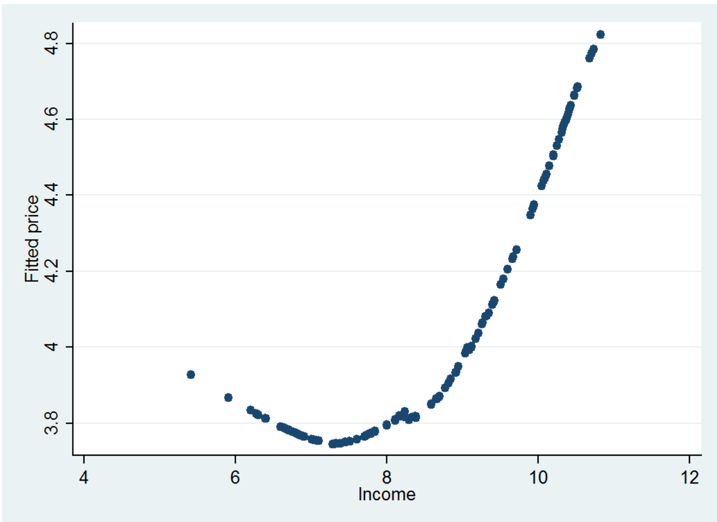

Figure 9.3: Penn World Table 8.0 (2005): non-parametric estimation of the price-income relation, fitted values