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Analyzing Food Price Trends in the Context of Engel's Law and the Prebisch-Singer Hypothesis

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Abstract

Income growth in emerging economies has often been cited as a key driver of the past decade's commodity price boom—the longest and broadest boom since World War II. This paper shows that in-come has a negative and highly significant effect on real food commodity prices, a finding that is consistent with Engel's Law and Kindleberger's thesis, the predecessors of the Prebisch-Singer hypothe-sis. The paper also shows that, in the long run, income influences real food prices mainly through the manufacturing price channel (the deflator), hence weakening the view that income growth exerts strong upward pressure on food prices. Other (short-term) drivers of food prices include energy costs, inventories, and monetary conditions.

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Even if one accepts the conclusion that the net barter terms have deteriorated for underdeveloped countries, there still remain the more significant questions of why this deterioration has occurred and what is the upshot for accelerating development in poor countries.

Gerald M. Meier (1958)

1. Introduction

After a nearly three-decade decline, world food commodity prices doubled in just six years, 2002-08. The price boom began as emerging economies entered an unprecedented growth period. GDP of low- and middle-income countries, for example, grew at an annual average of 6.4 percent during 2004-13, the highest of any 10-year period since 1960. Numerous commentators argued that the post-2005 upward trend of food commodity prices was due to the growing number of people in emerging economies, especially China and India, who are becoming wealthy enough to emulate Western diets (e.g., Krugman 2008; Wolf 2008). The relationship between income-induced demand growth and food commodity prices has been highlighted by a number of authors as well (Roberts and Schlenker 2013; Hochman, Rajagopal, and Zilberman 2010; Gilbert 2010).

Historically, this apparent link between income growth, food consumption, and food prices has not been expressed so pervasively, especially as a long-run relationship. As early as the mid-19th century, the German statistician Ernst Engel observed that poor families spend a greater proportion of their total expenditure on food compared to their wealthier counterparts, thus leading to the so-called Engel's Law of less than unitary income elasticity of food commodities. A century later, Kindleberger (1943, p. 349) argued that "… the terms of trade move against agricultural and raw material countries as the world's standard of living increases … and as Engel's Law of consumption operates." Kindleberger's thesis, which was verified by Prebisch (1950) and Kindleberger (1958) himself and also emphasized by Singer (1950), was later coined the Prebisch-Singer (PS) hypothesis. The declining path of the terms of trade (ToT) formed the intellectual foundation for post-WWII industrialization policies—taxation of primary commodity sectors in favor of manufacturing products, particularly in developing countries.

In view of the past decade's commodity price boom and strong income growth, this paper revisits the PS hypothesis. Using panel Autoregressive Distrusted Lag models, we show that income has a negative impact on the ToT of agricultural commodities, confirming the PS hypothesis. Moreover, the principal channel through which income impacts ToT is manufacturing prices, thus weakening the view that income growth by emerging economies has played a key role during the past decade's run up of food prices. We further show that in the short run, income positively affects both ToT and nominal commodity prices. Other drivers of food commodity prices include energy costs, physical stocks, and monetary conditions.

2. From Engel to Kindleberger and Prebisch-Singer

Based on expenditures of 153 Belgian families in 1853, Engel (1857) noted that "[t]he poorer a family, the greater the proportion of its total expenditure that must be devoted to the provision of food" and later concluded that "… the wealthier a nation, the smaller the proportion of food to total expenditure." (Quoted in Stigler 1954, p. 98). Following Engel's observation, numerous competing theories attempted to explain the long term behavior of the ToT faced by developing countries (Rostow 1950, Kindleberger 1958, Toye and Toye 2003). The first view focused on the supply-side and predicted that primary commodity prices will increase faster than manufacturing prices due to the resource constraints of the former and technological improvements of the latter—a view somewhat consistent with a Malthusian outcome. A second view assumed that ToT will follow investment cycles: investment expansion will induce a supply response in manufacturing goods, leading to lower prices, thus increasing the ToT. Conversely, an investment contraction would lead to a declining ToT. Proponents of a third view argued that ToT will follow a downward path because income growth leads to smaller demand increases in primary commodities than manufacturing products, a path consistent with Engel's Law.

While no dominant view emerged during the early part of the 20th century, the negative ToT-income relationship became the prevailing position after World War II. A turning point was Kindleberger's (1943) judgment that, in order for commodity-dependent countries to mimic the growth path of rich nations, they must transfer resources from agriculture and mining sectors to manufacturing industries. Other authors, however, warned against such interventions (see Akiyama and others (2003) for a literature review in the African context). Johnson (1947) argued that the agricultural sector should not be subjected to interventions. Friedman (1954) disputed the benefits of managing income variability for agricultural producers. Johnston and Mellor (1961) criticized the pro-urban policies pursued by many developing countries. Even Prebisch (1950) was cautious on how far industrialization policies should be pushed. Despite such warnings and cautious assessments, Kindleberger's view dominated the post-war development agenda and set the stage for the industrialization policies pursued by low-income countries.

Little attention was paid to the long-term ToT trend during the first few decades after World War II, most likely because commodity prices did not exhibit consistent trends or move in a synchronous manner. For example, while crude oil prices remained virtually unchanged and metal prices experienced sustained increases during the three decades after the war, agricultural prices on the whole declined, although subject to two spikes (during the Korean war in the early 1950s and just before the 1973 oil crisis). Interest in analyzing ToT reemerged in the 1980s for at least three reasons. First, following the 1970s boom, virtually all commodity prices declined, subsequently stabilizing at

much lower levels in real terms (figures 1 and 2). Therefore, the PS hypothesis began to fit the data well.

Second, the rationale and effectiveness of the bias against primary commodity sectors was being questioned. Bates (1981) called for a reconsideration of policies on commodity markets in order to promote economic development in rural communities. Lal (1985) criticized pricing policies and marketing arrangements in commodity-dependent developing countries. The tide took a definite turn against industrialization policies with the publication of two influential reports: the 1986 World Development Report, which highlighted the problems associated with policy interventions in agricultural commodity markets (World Bank 1986) and a detailed assessment of distortions associated with primary commodity sectors (Krueger, Schiff, and Valdès 1992).

Third, research on the long-term behavior of ToT was aided by two important papers. With regard to estimation, Engle and Granger (1987) introduced cointegration—a procedure that enabled researchers to separate meaningful long-term relationships from spurious correlations. On the subject of data, Grilli and Yang (1988) constructed a price index consisting of 24 internationally-traded primary commodities since 1900.

At least 45 studies have examined the PS hypothesis since the 1980s, half of which support the PS hypothesis (Appendix A). The mixed evidence is not surprising given that ToT followed different paths across commodity groups and time. Interestingly, the results are mixed even for the studies that used the Grilli-Yang data, implying that conclusions regarding the PS hypothesis are sensitive to methodology as well (of the 40 studies that tested the PS hypothesis since 1988, 23 have utilized the Grilli-Yang data).

3. Modeling Food Price Trends

The testable implications of the ToT-income relationship can be examined within a 2-sector, closed-economy framework as a move from equilibrium E_1 to E_2, in response to an exogenous income shock (figure 3). Let Q_i^F, Q_i^M and P_i^F, P_i^M denote consumption and prices of the primary and manufacturing commodities in period i, and Y_i denote income, i = 1, 2. Income level in the first period, Y_1, is consistent with consumption bundle $[Q_1^F, Q_1^M]$ and a price ratio of (P_1^F/P_1^M). As income increases to Y_2 in period 2, it leads to a consumption bundle of $[Q_2^F, Q_2^M]$ and a price ratio of (P_2^F/P_2^M). Now assume neutral technical change and non-homothetic preferences such that the increase in the consumption of the manufacturing commodity is larger than its primary commodity counterpart, $Q_2^M - Q_1^M > Q_2^F - Q_1^F$. These assumptions imply that $(P_2^F/P_2^M) < (P_1^F/P_1^M)$, hence the inverse ToT-income relationship consistent with Kindleberger's thesis and, by extension, Engel's Law. Conversely, under homothetic preferences and biased technical change against primary commodities, a positive ToT-income relationship will emerge, consistent with the Malthusian path discussed earlier. (Note that non-homotheticity of

preferences and biased technological change are not mutually exclusive since the former applies to the preference space while the latter applies to the production space.)

From an empirical perspective, the ToT-income relationship can be written as:

$$ToT_t = \beta_0 + \beta_1 Y_t + \varepsilon_t, \tag{1}$$

where subscript t denotes the continuous-time counterpart of the 2-period representation and $ToT_t = P_t/P_t^M$; β_1 is expected to be negative. Most of the literature on the PS hypothesis (see Appendix A), however, analyzed the long term behavior of ToT in the form of (2) or variations that allow for non-linear trends, structural breaks, or panel effects:[1]

$$ToT_t = \gamma_0 + \gamma_1 t + \varepsilon_t. \tag{2}$$

The PS hypothesis is confirmed if $\gamma_1 < 0$. Such specification, while informative, could mask the original intention of Kindleberger, as income may not necessarily grow in a linear fashion.

Unlike earlier studies, our empirical analysis directly tests for the long-term impact of income on ToT as originally envisaged by Kindleberger. Moreover, we extend the literature on PS hypothesis by asking why the ToT behaved in a certain way, not just if it did. This question was asked by Meier (1958, p. 88) when commenting on Kindleberger (1958): "Even if one accepts the conclusion that the net barter terms have deteriorated for underdeveloped countries, there still remain the more significant questions of why this deterioration has occurred and what is the upshot for accelerating development in poor countries." To this end, we include key sectoral and macroeconomic fundamentals in equation (1), thereby examining the validity of the PS hypothesis and the "Prebisch-Singer effect" by holding the other fundamentals constant. We further examine whether the income effect on ToT operates through primary commodity prices, manufacturing prices, or both, thus quantifying "Engel's Law effect."

We begin by rewriting (1) as

$$ToT_t = \beta_0 + \beta_1 Y_t + \mathrm{B}'\boldsymbol{F}[\bullet] + \varepsilon_t, \tag{3}$$

[1] Another change that took place between the early ToT debate and the empirical verification of the PS hypothesis relates to the measure of ToT. Earlier on, ToT was defined as the ratio of export to import prices. Most of the studies on the Prebisch-Singer hypothesis, however, defined ToT as the ratio of world primary commodity prices over the Manufacture Unit Value (MUV) index. Such a distinction has been highlighted by various authors, including Myint (1954-55) who noted that the two measures have vastly different implications for developing countries. From a policy perspective, this distinction is even more important today given that several emerging economies produce and export both primary and manufacturing commodities. Furthermore, a number of high income countries could fit into the commodity-dependent country profile (e.g. Australia, Canada, New Zealand.)

where $\boldsymbol{F}[\bullet]$ denotes the vector of sectoral and macroeconomic fundamentals and B is a vector of the respective coefficients. Equation (3) can be viewed as a reduced-form price-determination model, a result of equating aggregate demand and supply of a commodity, and subsequently expressing the equilibrium price as a function of quantifiable fundamentals. The theoretical underpinnings of the model are outlined in Holtham (1988) and Deaton and Laroque (1992). On empirical applications, Gilbert (1989) looked at the effect of developing countries' debt on commodity prices; Pindyck and Rotemberg (1990) examined comovement among various commodity prices; Reinhart (1991) and Borensztein and Reinhart (1994) analyzed the factors behind the weakness of commodity prices during the late 1980s and early 1990s; Frankel and Rose (2010) analyzed the effects of various macroeconomic variables on agricultural and mineral commodities; Baffes and Dennis (2015) analyzed the drivers of the post-2004 food price increases.

Expression $F[\bullet]$ is approximated by two macroeconomic indicators, US dollar interest rates (R_t) and exchange rates (X_t), and two sectoral fundamentals, stock-to-use ratios (S_t) and energy prices (P_t^E):

$$ToT_t = \beta_0^i + \beta_1^i Y_t + \beta_2^i R_t + \beta_3^i X_t + \beta_4^i S_{t-1}^i + \beta_5^i P_t^E + \varepsilon_t^i, \tag{4}$$

where i represents individual commodity; all variables are expressed in logarithms, and hence the $\widehat{\beta}_j^i$s can be interpreted as elasticities.

Testing the PS hypothesis requires interpreting the parameters of equation (4) in a corresponding manner. This implies that income growth will have a negative impact on the ToT. The same holds for interest rates since their low level is expected to reduce the cost of holding inventories, in turn leading to higher inventories and hence stronger demand. The depreciation of the U.S. dollar is consistent with higher commodity prices since it strengthens the demand and limits the supply from non-U.S. dollar consumers and producers. Low physical stocks relative to consumption (typically associated with tight supply conditions) lead to higher prices, and vice versa. To account for likely simultaneity between prices and stocks, the stock-to-use ratio is expressed in lagged form. Finally, higher energy prices, through their impact on production costs, are expected to be associated with higher commodity prices. Hence, the expected signs of the parameter estimates (noted as superscripts) are: β_1^-, β_2^-, β_3^-, β_4^-, β_5^+.

If the PS hypothesis holds ($\beta_1 < 0$), addressing the issue of why ToT has declined requires relaxation of the homogeneity assumption of manufacturing prices by re-parameterizing equation (4) using the fact that $ToT_t = P_t/P_t^M$:

$$P_t^i = \beta_0^i + \beta_1^i Y_t + \beta_2^i R_t + \beta_3^i X_t + \beta_4^i S_{t-1}^i + \beta_5^i P_t^E + \beta_6^i P_t^M + \varepsilon_t^i \tag{5}$$

Equation (5) describes how nominal food prices respond to income by controlling for sectoral and macroeconomic fundamentals as well as manufacturing prices. If income growth has led to an increase in food commodity prices, then $\beta_1 > 0$. Furthermore, if the

homogeneity restriction holds, i.e., $\beta_6 = 1$, equation (5) would be equivalent to (4).

Last, it should be noted that the primary purpose of this paper is to examine the ToT-income relationship and the channel through which the latter affects the former. We do not attempt to model the supply-side factors on commodity prices, which are likely to play an important role due to the nature of agricultural production. However, we do account for the impact of supply-side factors by including the stocks-to-use ratio of each commodity in equations (4) and (5).

4. Data

We consider six agricultural commodities, maize, soybeans, wheat, rice, palm oil, and cotton. The inclusion of cotton was motivated by the desire to account for as much of world's arable land as possible. Commodity prices represent annual (calendar) averages, expressed in U.S. dollar per metric ton for crops and in US dollars per barrel for crude oil, all of which are obtained from the World Bank's Commodity Price Database.

We use the manufacture unit value (MUV) as a proxy for manufacturing prices. The MUV is a US dollar trade-weighted index of manufactures exported from 15 economies (Brazil, Canada, China, Germany, France, India, Italy, Japan, Mexico, the Republic of Korea, South Africa, Spain, Thailand, the United Kingdom, and the United States). The MUV has been used extensively in the literature as a deflator of nominal commodity prices when calculating real prices, or ToT (almost all studies reported in Appendix A that have used the Grilli-Young data have utilized MUV).[2]

Given the study's objective to identify the effect of income on food prices, 16 alternative income measures were used, by measuring GDP at market prices and Purchasing Power Parity (PPP) terms for four aggregation levels following the World Bank's classification (world, high-, middle-, and low-income countries), both in global and per capita terms (figure 4). For the real interest rate, the 3-month U.S. Treasury bill, adjusted by the U.S. Consumer Price Index was used. The exchange rate was represented by the U.S. dollar Real Effective Exchange Rate against a broad basket of currencies. The stocks-to-use ratio represents the ratio of global ending stocks to global consumption at the end of the marketing year, obtained from the U.S. Department of Agriculture's

[2] The description of commodity prices is as follows (country names in parentheses indicate where the transaction takes place; they do not correspond to the average prices received by producers or paid by consumers): **maize** (United States), no. 2, yellow, f.o.b. (free on board) U.S. Gulf ports; **rice** (Thailand), 5 percent broken, white rice, milled, indicative price based on weekly surveys of export transactions, government standard, f.o.b. Bangkok; **wheat** (United States), no. 1, hard red winter, ordinary protein, export price delivered at the U.S. Gulf port for prompt or 30 days shipment; soybeans (United States), c.i.f. (cost, insurance, freight) Rotterdam; **palm oil** (Malaysia), 5 percent bulk, c.i.f. N. W. Europe; **cotton** (Cotton Outlook "Cotlook A index"), middling 1-3/32 inch, traded in Far East, C/F beginning 2006; previously Northern Europe, c.i.f.; and crude oil, average price of Brent, Dubai and West Texas Intermediate, equally weighed. Information of price data and MUV can be found at http://worldbank.org /commodities

Production, Supply and Distribution Online database. The sample covers the 1960-2014 period yielding 55 annual observations.

5. Estimation Procedure

The nature of our model—a long run relationship applied to 55 annual observations for six commodities—can be represented best within a panel framework. However, the likely presence of common shocks and spatial dependence (common features of commodity prices) implies that panel estimation procedures assuming no cross-sectional correlations are likely to be inconsistent. Therefore, we first use the Pesaran (2004) approach under both a standard fixed and random effects framework, as well as the Breusch-Pagan test under a fixed effect model to test for cross-sectional dependence. As can be seen in table 1, the null hypothesis of cross-sectional independence is overwhelmingly rejected at 1 percent significance level.

A second concern is the stationarity properties of variables that vary by commodity (ToT, nominal prices, and stocks-to-use ratios). To address this concern, we use the panel unit root test of Pesaran (2003) which is based on the mean of individual augmented Dickey-Fuller (ADF) t-statistics of each panel.[3] Elimination of cross-sectional dependence is achieved by augmenting the standard ADF regressions with the cross-sectional averages of lagged levels and first differences of the individual series. Results in table 2 confirm rejection of the null hypothesis of no unit root for ToT and nominal prices at most lags in all panels, both with and without a time trend. By contrast, the stocks-to-use ratio does not contain a unit root when no trend is considered in the estimation equation.

For the variables that do not vary by commodity, we test for the presence of unit roots by using the ADF and the Phillips-Perron (PP) tests (Phillips and Perron 1988) with and without trend. Results reported in table 3 show that the unit root null hypothesis cannot be rejected for most of the income measures when a trend is included. In particular, the income is stationary when measured by GDP of high-income countries, regardless of the test used or the indicator considered. Nevertheless, the overall picture is that the level of income tends to contain a unit root, in favor of first-difference stationary. Similar conclusions apply to the exchange rate, interest rate, manufacturing prices, and oil prices.

We next test for the presence of a cointegration relationships between ToT and income. As was the case in the panel unit root tests, the first generation of panel cointegration tests relied on the assumptions of homogeneity and independence across

[3] The first generation of panel unit root tests, such as those proposed by Levin et al. (2002) and Im et al. (2003) assumed cross-sectional independence. See Gengenbach et al. (2010) for a recent review on panel unit root tests.

panels. In the context of our model we consider the approach proposed by Westerlund (2007) which is robust to not only heterogeneity in the long-run cointegrating relationship and short-run dynamics, but also to dependence in both cross-sectional and time dimensions. Results from the panel cointegration tests are shown in table 4. We report two versions of the panel test—one uses standard errors and the other uses the Newey-West adjusted standard errors allowing for heteroskedasticity and autocorrelation. With few exceptions, the null hypothesis of no cointegration relationship between prices and income cannot be rejected at the 10 percent significance level.

Last, we use a panel Autoregressive Distributed Lags (ARDL) model to estimate the long- and short-run impacts of income on ToT and nominal prices. In a series of papers, Pesaran and colleagues show that the ARDL estimator can be used for long-run analysis even in the presence of reverse causality and nonstationary regressors (Pesaran and Smith 1995; Pesaran 1997; and Pesaran and Shin 1999).[4] The panel ARDL approach was further extended in various studies, including Pesaran (2006), Kapetanios, Pesaran, and Yamagata (2011), Cesifo et al. (2014).[5]

Consider the following ARDL specification of the ToT model under a panel framework:

$$ToT_t^i = \beta_0^i + \sum_{p=1}^{P} \lambda_p^i ToT_{t-p}^i + \sum_{k=0}^{K} \beta_{1k}^i Y_{t-k} + \mathrm{B}'\boldsymbol{F}[\bullet] + \varepsilon_t^i, \tag{6}$$

where $p = 1,2, \ldots . P$ is the lag order of ToT, and $k = 1,2, \ldots . K$ is the lag order of the income measure, λ_p^i is the coefficient associated with each autoregressive term. Since our primary purpose is to examine the impact of income on ToT, for simplicity and parsimony we do not include any lags of the fundamental factors $\boldsymbol{F}[\bullet]$. The panel ARDL(P,K) model in its error-correction form can be written as:

$$\Delta ToT_t^i = \beta_0^i + \delta_i\left(ToT_{t-1}^i - \theta_i Y_t\right) + \sum_{p=1}^{P-1} \lambda_p^{i*} \Delta ToT_{t-p}^i + \sum_{k=0}^{K-1} \beta_{1k}^{i*} \Delta Y_{t-k} + \mathrm{B}'\boldsymbol{F}[\bullet] + \varepsilon_t^i, \tag{7}$$

where $\delta_i = -(1 - \sum_{p=1}^{P} \lambda_p^i)$ is the speed of adjustment to the long-run equilibrium

[4] Other commonly used tests include the dynamic least square estimator of Kao and Chiang (2000) and the continuously-updated and bias-corrected estimator of Bai, Kao, and Ng (2009). However, these approaches may not be appropriate in the present context given the likelihood of reverse causality in some of the income measures—for countries relying on commodity exports, an increase in commodity prices may result in higher income. Moreover, the DOLS approach assumes cross-sectional independence among error terms, an assumption that does not hold for commodity prices as noted earlier.

[5] In Pesaran (2006) the panel ARDL model is estimated in an augmented version by including the average of dependent variable and independent variable across panels as regressors. However, in the current study the income measure is the same across panels, and hence including these terms will lead to collinearity.

between income and ToT, $\theta_i = (\sum_{k=0}^{K} \beta_{1k}^{i})/(1 - \sum_{p=1}^{P} \lambda_p^i)$ is the long-run effects of income on ToT, $\lambda_p^{i*} = \sum_{m=p+1}^{P} \lambda_m^i$ for $p = 1,2, \ldots, P-1$, and $\beta_{1k}^{i*} = \sum_{h=k+1}^{K} \lambda_h^i$ for $k = 1,2, \ldots, K-1$. Both the speed of adjustment and long-run coefficients can be calculated by replacing the parameters with the short-run analysis from (6). Note that equation (7) allows for a significant degree of heterogeneity across panels as coefficient estimates can vary among different cross-sectional units.

A consistent estimate of the average panel long-run effect can be obtained from individual estimates across panel units. First, we consider the mean-group (MG) estimator that allows intercepts, slope coefficients, and error variances to differ across cross-sectional units. We also use the panel mean-group estimator (PMG) that constrains the long-run effect to be the same across panels but allows for heterogeneous short-run estimates as well as the speed of adjustment coefficient for different cross-sectional units.

Because the long-run coefficient (θ_i) is dependent on the estimates of lagged dependent variable terms, even small changes in λ could induce large changes on θ. Hence, for robustness we use two lag lengths. For simplicity, the same lag length is used for both ToT and income. Results of the long-run effect of income on ToT as well as the speed of adjustment of ToT to this long-run relationship are reported in table 5.

6. Discussion

6a. The Role of Income

As can be seen from table 5, with only a few exceptions (seven out of 64 cases), income has a negative and statistically significant impact on ToT in the long-run, thus rendering strong support for the PS hypothesis. These results are robust to the lag length and type of estimator considered. All non-significant long-run income parameter estimates are associated with per capita GDP in low-income countries, an indication of weak "Prebisch-Singer" and "Engel's Law" effects for this income measure. However, when the total GDP of low-income countries used, the PS hypothesis continues to hold.

Beyond confirming the PS hypothesis, the size of elasticities from various income measures is of interest in its own right. To put the income effect into perspective, consider that in the two decades spanning 1994-2003 and 2004-13, world per capita GDP measured at PPP increased by 36 percent (logarithmic changes). Applying the corresponding income elasticity of -0.79 (parameter estimate from one lag) implies that the "Prebisch-Singer effect" reduced the ToT by 28 percentage points [-28.44% = -0.79*36%]. Stated otherwise, if income did not change, the ToT would have almost doubled during these two decades (an increase of 64 percent instead of 36 percent).

For all estimates that are significantly different from zero, the long-run income elasticity ranges from -1.05 to -0.34. When measured in per capita terms, GDP of high income countries exhibits a larger elasticity than middle-income countries, which is in turn larger than the elasticity for low-income countries. As expected, elasticities based on

per capita income are higher than their total income counterparts due to the latter's dampening effect from population growth. With some exceptions, the sizes of these elasticities are inversely related to the income growth rates observed during the sample period, implying that when the elasticity estimates are combined with income growth, the effect of income on ToT is similar across income measures. Finally, the results reported in table 5 suggest that when deviations to the long-run equilibrium between ToT and income occurs, ToT adjusts with rates rate of 33 to 58 percent annually. All the adjustment coefficients are statistically significant at the 1 percent level.

Yet it is not clear from these estimates whether income's negative impact on ToT operates through nominal commodity prices, manufacturing prices, or both. It is possible that income may have a positive effect on primary commodity prices, but this effect is not sufficient to induce a positive long-run income impact on ToT due to the faster rate of growth in manufacturing prices. Equivalently, income may not have a positive impact on nominal prices in the long-run, giving rise to negative income impact on ToT. In order to identify the transmission mechanism, we estimate panel ARDL models that examine the long-run equilibrium among nominal commodity prices, income, and manufacturing prices, the results of which, based on the MG estimator, are reported in table 6. Parameter estimates based on PMG procedure are similar.

With only one exception, the parameter estimate of manufacturing prices is positive and significantly different from zero, a result which holds all income groups and lag lengths. The only exception is total GDP measured at market price of middle-income countries and a lag length of two. However, the magnitude of the impact varies considerably, with the estimated price elasticity ranging from 0.19 to 0.59. Using the mean of these estimates (0.42), an increase of manufacturing prices by 10 percent in the long-run is associated with a 4.2 percent increase in nominal commodity prices—less than half of what the homogeneity restriction would necessitate. The mean of estimated elasticities of manufacturing prices against nominal prices is 0.43, 0.48, 0.28, and 0.32 for GDP of world, high-income, middle-income, and low-income countries, respectively.

The results for the long-run impact of income on nominal prices, however, are at best mixed. Out of 32 estimates (16 income measures, two lag lengths), only four are significantly positive, all of which are associated with per capita income in low- and middle-income countries. Income does not affect nominal commodity prices in in any significant way in the remaining 28 cases. The only exceptions are the per capita income measures of low and middle-income countries, both of which have a positive long-run effect on nominal agricultural prices, with an estimated income elasticity ranging from 0.27 to 0.56. Overall, results from the panel ARDL estimation of nominal prices suggest that income does not have a significant impact on nominal prices.

Given that most of the elasticity estimates are not statistically significant, the role of income growth on primary commodity prices appears to be less important than typically assumed. Even when measured with the GDP of middle- and low-income

countries, the evidence is not conclusive. Moreover, the parameter estimate of manufacturing prices suggests that in the long-run, the response of nominal prices to manufacturing prices is far lower than unity. These findings indicate that manufacturing prices may be the key (and, perhaps, the only) transmission channel through which income affects ToT in the long-run.

Although these results are somewhat different from the widely held belief that income-driven demand growth has exerted a strong influence on food commodity prices, in some ways they should be expected. The empirical literature on the subject is, at best, mixed. Pindyck and Rotemberg (1990) concluded that industrial production—a demand proxy—did not affect prices in any significant way when using a sample of seven commodities and more than two decades of monthly data. A similar conclusion is reached by Ai, Chatrath, and Song (2006), who consider U.S. CPI-deflated quarterly prices for five commodities. More recently, Frankel (2014) applied OLS regressions and a panel specification on annual U.S. GDP-deflated prices for 11 primary commodities and found that overall GDP did not affect commodity prices.

Nonetheless, the finding of a weak income-food price relationship is not shared by all. Gilbert (2010), for example, concluded that demand-side variables, including GDP growth, are important in explaining aggregate food price indices. Hochman et al. (2011) found that income accounted for almost one third of 2001-08 food price increases. Erten and Ocampo (2013) investigated the presence of super cycles during 1865-2010 and found evidence in favor of the Prebisch-Singer hypothesis for non-oil commodities.

On the other hand, evidence that grain consumption by emerging economies has experienced growth rates that are either high by historical standards or comparable to their income growth rates is, at best, weak. Alexandratos (2008) concluded that China's and India's combined average annual increase in grain consumption was lower in 2002-08 than in 1995-2001. Other studies have reported similar findings, including Alexandratos and Bruinsma (2012), Sarris (2010), Baffes and Haniotis (2010), FAO (2009), and Lustig (2008). In fact, Deaton and Drèze (2008) found that, despite growing incomes, there has been a downward trend in calorie intake in India since the early 1990s.

Furthermore, it is important to emphasize that the results on Engle's Law, PS hypothesis and observed consumption patterns for food commodities do not necessarily hold for industrial commodities. Webster, Paltsev, and Reilly (2008), for example, based on a review of the parameter values used in various integrated assessment models reported that the income elasticity for energy demand exceeds unity. The role of income-driven demand growth (and hence price increases) in industrial commodities has been confirmed by numerous authors; see, for example, Kilian (2009) for the oil market; Baffes (1997), Labys, Achouch, and Terraza (1999), and Issler, Rodrigues, and Burjack (2014) for base metals. These results are consistent with consumption patterns of emerging economies, especially China which now accounts for nearly half of the world's metal consumption, up from a mere 4 percent two decades ago, and non-OECD economies,

which currently consume more than half of world's crude oil, up from one third 15 years ago.[6]

Finally, Table 7 presents the short-run parameter estimates from the panel ARDL regression models with income represented by GDP per capita measured at PPP for both ToT and nominal prices using the MG estimator (equation 7). As can be seen, income overall has a positive significant impact on both ToT and nominal prices in the short-run.

6b. The Role of Other Fundamentals

The discussion thus far focused on the long-run impact of income. However, it is the remaining fundamentals, as represented by $F[\bullet]$, that may play a key role in the short-run food price determination process. As can be seen from table 7, the stock-to-use ratio estimates are negative and highly significant across all estimates, ranging from -0.44 to -0.32 for the ToT model and from -0.36 to -0.29 for the nominal price model. These elasticity estimates are comparable to figures reported elsewhere. For example, Bobenrieth, Wright, and Zeng (2013) estimated correlation coefficients between stock-to-use ratios and real de-trended prices for maize, wheat, and rice of -0.50, -0.40, and -0.17, respectively. Similarly, FAO (2008, p. 6) reported correlation coefficients between the cereals price index and various measures of stock-to-use ratios in the range of -0.47 to -0.65.

With a few exceptions, the estimate of the oil price elasticity was positive and significantly different from zero across different income measures and lag lengths for both ToT and nominal price models. The elasticity estimates averaged 0.07 across the ToT models and 0.15 across the nominal price models (when the significant parameter estimates are considered).

The relationship between energy and non-energy prices has been established long before the recent price boom. Gilbert (1989) estimated transmission elasticity from energy to non-energy commodities of 0.12 and from energy to food commodities of 0.25. Hanson, Robinson, and Schluter (1993) based on a CGE model found a significant effect of oil price changes to agricultural producer prices in the United States. Borensztein and Reinhart (1994), estimated transmission elasticity to non-energy commodities of 0.11. A strong relationship between energy and non-energy prices was found by Chaudhuri (2001) as well. Baffes (2007), estimated elasticities of 0.16 and 0.18 for non-energy and food commodities, respectively. Moss, Livanis, and Schmitz (2010) found that U.S. agriculture's energy demand is more sensitive to price changes than any other input.

Yet not all studies concur with a strong oil/non-oil price relationship. Saghaian (2010) established strong correlation among oil and other commodity prices (including

[6] To the extent that strong growth by emerging economies boosts energy demand and hence oil prices, food prices have been driven in part by the growth patterns of these economies. In fact, this point has been highlighted by Heady and Fan (2010) and Baffes and Dennis (2015).

food) but the evidence for a causal link between oil and other commodities was mixed. Gilbert (2010) established a correlation between crude oil and food prices both in terms of levels and changes, but noted that it could reflect common causation, not a causal link. Zhang and others (2010) found no direct long-run relationship between fuel and agricultural commodity prices and a weak short-run relationship. Reboredo (2012) concluded that the prices of maize, wheat, and soybeans were not driven by oil price fluctuations.

The mixed evidence on the energy/non-energy price link may reflect the expansion of biofuel use during the past decade. Consider, for example, that exogenous shocks pushing crude oil prices up (down) would lower (increase) fuel consumption, thus "delinking" the prices of oil from those of biofuel commodities in the presence of mandated biofuel consumption (de Gorter and Just 2009). The mixed evidence could also reflect the frequency of the data—higher frequency (and hence "noisier") data are typically associated with weaker correlations (Zilberman et al. 2013).

Parameter estimates of real exchange rates in both ToT and nominal price models are all negative and the majority are statistically significant as well. The magnitude of elasticity ranges from -0.39 to -0.35 for ToT and from -0.71 to -0.56 for nominal prices. This is highly consistent with expectations, as this effect also reflects the relative share of US agricultural production and trade in these markets (Lamm 1981; Gardner 1981).

Although interest rates have a negative impact on ToT, it is statistically different from zero in only one case. For the nominal price model, interest rates have a positive impact on prices in five out of eight cases. However, the magnitude of such impact is small, with the elasticity estimate ranging from 0.01 to 0.02. The mixed impact of interest rate on commodity prices is consistent with the empirical literature. Gilbert (1989) concluded that high interest rates have a negative impact on the metal price index, though with considerable lags. Baffes (1997) estimated mostly negative, yet not significantly different from zero, elasticity for five metal prices. Akram (2009) concluded that commodity prices increase significantly in response to a reduction in real interest rates. Frankel and Rose (2010) found little to support that easy monetary policy exerts upward pressure on real commodity prices. In a subsequent paper, Frankel (2013) focused on the trade-off between interest rates and expectations on future price changes (as a speculative factor) but concluded that in only one out of seven cases of the model's panel version, the effect of interest rates on commodity prices was significantly different from zero. Anzuini, Lombardi, and Pagano (2013) did establish that easy monetary policy is associated with higher commodity prices, but also noted that the impact is modest. Byrne, Fazio, and Fiess (2013) found a significant negative relationship between real interest rates and real commodity prices, with shocks in real interest rates being absorbed in commodity prices within a five-year window. Last, Baffes and Savescu (2014) found mixed impact of nominal short term interest rates on six metal prices but also concluded that longer-term rates positively affect metal prices.

7. Conclusion

This paper reconciled the Prebisch-Singer hypothesis with the popular view that income growth in emerging economies has been the key driver of food price movements. In particular, the paper extended the literature on the income-food price relationship in three ways. First, it tested the hypothesis by examining the effect of income on the ToT, as originally envisaged by Kindleberger, rather than relating ToT with a time trend, which is the common practice in the literature. Second, in addition to income, it accounted for key sectoral and macroeconomic fundamentals that are expected to influence commodity prices. Third, it examined whether the primary channel of the effect of income on ToT is food commodity prices or manufacturing prices.

The paper employed a reduced-form price determination model, applied to 1960-2014 annual data for five food commodities (maize, soybeans, wheat, rice, palm oil) and cotton. Using panel Autoregressive Distributed Models, we show that income has a negative and highly significant effect on the terms of trade in the long run. This finding is consistent with both Engel's Law and Kindleberger's thesis, the predecessor of the Prebisch-Singer hypothesis. Moreover, results indicate that income's impact operates mainly through the manufacturing price channel (the deflator).

Other key drivers of food prices in the short run are energy costs, physical stocks, and monetary conditions. Crude oil and crop conditions are important drivers as well. A depreciating U.S. dollar positively affects nominal food prices and ToT. Finally, the effect of interest rates on the ToT are muted.

From a methodological perspective, the literature review revealed that the research on the Prebisch-Singer hypothesis applies mostly to the behavior of primary commodity prices, not manufacturing prices. In view of the differences in the way food and industrial commodities respond to fundamentals, as well as the fact that the impact of income on the ToT of food commodity prices operates primarily through the manufacturing price channel, there is a need to broaden the scope of analytical approaches on commodity price developments to more accurately reflect their drivers. Such approaches should at least focus on the relationship between income and manufacturing prices as well as measurement aspects of the latter.

Table 1
Cross-Sectional Dependence Tests

	---------- Pesaran Method ----------				----- Breusch-Pagan Method -----	
	---------- ToT ----------		------- Nominal Prices ------		ToT	Nominal Prices
	Fixed	Random	Fixed	Random		
GDP measured at Market Prices, constant US$ 2010, total						
World	14.08***	13.95***	11.45***	11.35***	224.5***	179.3***
High Income	13.48***	13.35***	11.46***	11.36***	205.8***	175.9***
Middle Income	15.02***	14.87***	11.36***	11.27***	254.6***	183.3***
Low Income	15.68***	15.52***	11.38***	11.29***	275.6***	182.9***
GDP measured at Purchasing Power Parity (PPP), total						
World	14.46***	14.32***	11.44***	11.34***	236.5***	180.8***
High Income	13.58***	13.45***	11.46***	11.36***	208.2***	176.8***
Middle Income	15.04***	14.89***	11.36***	11.27***	255.3***	182.8***
Low Income	15.63***	15.47***	11.39***	11.3***	273.8***	182.5***
GDP measured at Market Prices, constant US$ 2010, per capita						
World	14.46***	14.32***	11.44***	11.35***	233.8***	181.8***
High Income	13.36***	13.23***	11.46***	11.36***	201.6***	174.8***
Middle Income	15.71***	15.55***	11.28***	11.19***	278.1***	185.5***
Low Income	16.78***	16.62***	11.35***	11.25***	315.6***	186.9***
GDP measured at Purchasing Power Parity (PPP), per capita						
World	15.02***	14.87***	11.43***	11.34***	253.8***	183.3***
High Income	13.50***	13.37***	11.45***	11.35***	204.9***	176.0***
Middle Income	15.68***	15.52***	11.29***	11.20***	276.9***	184.8***
Low Income	16.71***	16.56***	11.36***	11.26***	313.4***	186.5***

Notes: The null hypothesis is cross-sectional independence. The test statistics are based on Pesaran (2004) and Breusch-Pagan cross-sectional dependence tests under either a standard fixed effects or random effects model. One (*), two (**), and three (***) asterisks indicate significance at 10%, 5%, and 1% levels, respectively.

Table 2
Panel Unit Root Test Results on Terms of Trade (ToT), Nominal Prices, and Stocks-to-Use Ratio

	Lag Length			
	2	4	6	8
Terms of Trade (ToT)				
Without trend	-4.42***	-1.43*	-0.41	0.01
With trend	-4.34***	-1.07	0.42	0.25
Nominal Price				
Without trend	-2.94***	-0.66	0.50	0.35
With trend	-3.47***	-0.71	0.36	-0.02
Stocks-to-Use Ratio				
Without trend	29.98***	25.16**	19.05*	27.61**
With trend	31.02***	11.98	6.38	14.23

Notes: The null hypothesis is that no unit root exists in any of the panels. One (*), two (**), and three (***) asterisks indicate significance at 10%, 5%, and 1% levels, respectively.

Table 3
Unit Root Test Results for Various Income Measures, Exchange rate, Interest Rate, Oil Price, and MUV

	---------- Levels ----------				---------- First Difference ----------			
	--- No Trend ---		--- With Trend ---		--- No Trend ---		--- With Trend ---	
	ADF	PP	ADF	PP	ADF	PP	ADF	PP
GDP measured at Market Prices, Constant US$ 2010, total								
World	-3.19**	-4.37***	-3.04	-2.85	-3.83***	-4.25***	-5.01***	-5.30***
High income	-3.90***	-6.10***	-2.14	-1.99	-3.49***	-3.87***	-5.34***	-5.50***
Middle income	-0.60	-0.03	-2.15	-1.32	-3.24**	-3.66***	-3.21*	-3.61**
Low income	3.37	2.89	0.97	1.66	-2.60*	-3.49***	-4.04***	-4.11***
GDP measured at Purchasing Power Parity (PPP), total								
World	-2.19	-2.56	-3.01	-2.59	-3.82***	-4.32***	-4.31***	-4.82***
High income	-3.75***	-5.65***	-2.38	-2.24	-3.46***	-3.92***	-5.07***	-5.39***
Middle income	-0.42	0.40	-2.05	-1.11	-3.53***	-3.73***	-3.48**	-3.64**
Low income	3.41	2.98	0.99	1.73	-2.54	-3.30**	-4.03***	-4.00***
GDP measured at Market Prices, Constant US$ 2010, per capita								
World	-2.37	-3.05**	-3.20*	-3.12	-4.47***	-4.81***	-4.95***	-5.24***
High income	-3.55***	-4.98***	-2.04	-1.69	-3.82***	-4.27***	-5.37***	-5.55***
Middle income	0.32	1.04	-1.26	-0.54	-3.03**	-3.39**	-3.07	-3.41**
Low income	2.88	3.63	0.97	1.80	-2.40	-3.19**	-3.91**	-3.87**
GDP measured at Purchasing Power Parity (PPP), per capita								
World	-1.21	-1.39	-2.43	-2.21	-4.18***	-4.65***	-4.23***	-4.73***
High income	-3.41**	-4.63***	-2.28	-1.96	-3.78***	-4.31***	-5.12***	-5.45***
Middle income	0.54	1.58	-1.04	-0.21	-3.16**	-3.29**	-3.26*	-3.35*
Low income	3.04	3.79	0.97	1.87	-2.33	-3.00**	-3.89**	-3.77**
Other Variables								
Exchange Rate	-3.07**	-2.30	-3.32*	-2.53	-4.17***	-4.52***	-4.13***	-4.48***
Interest Rate	-2.30	-2.66*	-2.55	-2.80	-6.36***	-8.15***	-6.32***	-8.09***
MUV	-1.58	-1.53	-1.29	-0.96	-3.78***	-4.09***	-4.01***	-4.24***
Oil Price	-0.97	-0.98	-1.85	-2.04	-5.13***	-7.44***	-5.07***	-7.37***

Notes: The lag length of the ADF test is selected by the BIC criterion. One (*), two (**), and three (***) asterisks indicate significance at 10%, 5%, and 1% levels, respectively.

Table 4
Panel Cointegration Test Statistics

	---------- Terms of Trade (ToT) ----------				---------- Nominal Prices ----------			
	------- No Trend -------		------ With Trend -----		-------- No Trend ------		------ With Trend ----	
	P_t	P_a	P_t	P_a	P_t	P_a	P_t	P_a
GDP measured at Market Prices, constant US$ 2010, total								
World	-6.00***	-4.95	-6.13	-4.71	-5.60*	-4.21	-7.26**	-2.98
High income	-6.01***	-7.72**	-6.72**	-6.94	-6.20**	-4.01	-8.52***	-4.48
Middle income	-7.52***	-4.67	-9.86***	-6.71	-8.20***	-10.03**	-6.32	-9.72
Low income	-6.45***	-6.02	-10.75***	-20.66***	-8.43***	-10.82**	-6.79*	-13.92
GDP measured at Purchasing Power Parity (PPP), total								
World	-6.57***	-4.21	-7.51***	-3.93	-5.68*	-4.70	-5.39	-2.60
High income	-6.10***	-6.47	-6.32*	-8.26	-5.93**	-4.27	-7.86***	-4.31
Middle income	-4.85*	-2.54	-10.23***	-7.46	-8.12***	-9.20*	-7.53**	-21.14***
Low income	-6.39***	-5.96	-10.71***	-20.83***	-8.44***	-11.11**	-6.86*	-14.61*
GDP measured at Market Prices, constant US$ 2010, per capita								
World	-6.28***	-6.62*	-6.84**	-6.64	-5.75*	-7.99	-6.17	-4.60
High income	-6.23***	-8.21**	-6.81**	-7.94	-6.39**	-5.08	-8.41***	-6.12
Middle income	-7.10***	-5.87	-11.89***	-15.99***	-8.54***	-13.45***	-8.23***	-17.38***
Low income	-5.33**	-6.68*	-10.75***	-22.83***	-9.53***	-15.21***	-6.99*	-13.95
GDP measured at Purchasing Power Parity (PPP), per capita								
World	-6.72***	-5.89	-8.80***	-7.53	-6.11**	-8.50	-5.05	-5.76
High income	-6.20***	-7.62**	-6.28	-8.65	-6.10**	-5.33	-7.76***	-5.72
Middle income	-4.48	-3.18	-12.37***	-18.3***	-8.38***	-12.16***	-8.22***	-17.41***
Low income	-5.38**	-6.67*	-10.74***	-22.81***	-9.55***	-15.14***	-7.08*	-14.29*

Notes: P_t refers to the test statistics calculated using standard errors, and P_a refers to the test statistics calculated using Newey-West standard errors. In the ToT model, the cointegration relationship between ToT and income is examined. In the nominal price model, the cointegration relationship among nominal prices, income, and manufacturing prices are examined. One (*), two (**), and three (***) asterisks indicate significance at 10%, 5%, and 1% levels, respectively.

Table 5
Long-Run Estimates from Panel Autoregressive Distributed Lag Model, Terms of Trade (ToT)

	---------- Mean-Group (MG) Estimator ----------				----- Pooled-Mean Group (PMG) Estimator -----			
	------ Lag 1 ----------		--------- Lag 2 ------		-------- Lag 1 ----------		--------- Lag 2 ---------	
	θ	**λ**	**θ**	**λ**	**θ**	**λ**	**θ**	**λ**
GDP measured at Market Prices, constant US$2000, total								
World	-0.41***	-0.47***	-0.45***	-0.51***	-0.46***	-0.47***	-0.47***	-0.50***
High income	-0.43***	-0.48***	-0.49***	-0.53***	-0.51***	-0.49***	-0.53***	-0.53***
Middle income	-0.35***	-0.44***	-0.35***	-0.46***	-0.37***	-0.43***	-0.37***	-0.46***
Low income	-0.49***	-0.42***	-0.40***	-0.43***	-0.52***	-0.43***	-0.51***	-0.45***
GDP measured at Purchasing Power Parity, total								
World	-0.41***	-0.46***	-0.44***	-0.49***	-0.45***	-0.46***	-0.46***	-0.49***
High income	-0.43***	-0.48***	-0.50***	-0.53***	-0.52***	-0.49***	-0.54***	-0.53***
Middle income	-0.34***	-0.43***	-0.34***	-0.45***	-0.36***	-0.43***	-0.35***	-0.46***
Low income	-0.47***	-0.42***	-0.39***	-0.43***	-0.50***	-0.43***	-0.49***	-0.45***
GDP measured at Market Prices, constant US$2000, per capita								
World	-0.84***	-0.45***	-0.97***	-0.48***	-1.01***	-0.45***	-1.05***	-0.48***
High income	-0.60***	-0.49***	-0.69***	-0.54***	-0.73***	-0.50***	-0.75***	-0.54***
Middle income	-0.56***	-0.39***	-0.51***	-0.40***	-0.61***	-0.39***	-0.60***	-0.41***
Low income	-0.35	-0.32***	0.09	-0.34***	-0.55	-0.33***	-0.32	-0.34***
GDP measured at Purchasing Power Parity, per capita								
World	-0.79***	-0.43***	-0.87***	-0.45***	-0.91***	-0.43***	-0.94***	-0.46***
High income	-0.60***	-0.49***	-0.70***	-0.53***	-0.74***	-0.49***	-0.76***	-0.54***
Middle income	-0.55***	-0.39***	-0.47***	-0.39***	-0.58***	-0.39***	-0.56***	-0.40***
Low income	-0.37	-0.33***	0.10	-0.34***	-0.61*	-0.33***	-0.43	-0.34***

Notes: θ denotes the long-run impact of income on ToT, λ is the speed of adjustment to long-run equilibrium. One (*), two (**), and three (***) asterisks indicate significance at 10%, 5%, and 1% levels, respectively.

Table 6
Long-Run Estimates from Panel Autoregressive Distributed Lag Model, Nominal Prices, MPG Estimator

	---------- Lag 1 ----------			---------- Lag 2 ----------		
	---- θ_1 ---	---- θ_2 ----	---- λ ----	---- θ_1 ---	---- θ_2 ----	---- λ ----
GDP measured at Market Prices, constant US$ 2000, total						
World	-0.07	0.55***	-0.52***	-0.03	0.46***	-0.57***
High income	-0.08	0.55***	-0.53***	-0.06	0.49***	-0.58***
Middle income	0.04	0.42***	-0.52***	0.15	0.21	-0.58***
Low income	-0.02	0.46***	-0.53***	0.19	0.22**	-0.59***
GDP measured at Purchasing Power Parity, total						
World	-0.09	0.57***	-0.51***	-0.03	0.46***	-0.56***
High income	-0.10	0.58***	-0.53***	-0.07	0.50***	-0.58***
Middle income	0.03	0.44***	-0.52***	0.13	0.24*	-0.57***
Low income	0.00	0.46***	-0.53***	0.21	0.20**	-0.59***
GDP measured at Market Prices, constant US$ 2000, per capita						
World	-0.12	0.52***	-0.52***	-0.021	0.41***	-0.57***
High income	-0.14	0.57***	-0.54***	-0.14	0.52***	-0.59***
Middle income	0.12	0.40***	-0.53***	0.32**	0.19**	-0.59***
Low income	0.09	0.44***	-0.54***	0.56***	0.33***	-0.62***
GDP measured at Purchasing Power Parity, per capita						
World	-0.10	0.52***	-0.52***	0.051	0.36***	-0.57***
High income	-0.17	0.59***	-0.53***	-0.15	0.53***	-0.58***
Middle income	0.09	0.42***	-0.52***	0.27*	0.22**	-0.58***
Low income	0.11	0.44***	-0.54***	0.56***	0.32***	-0.63***

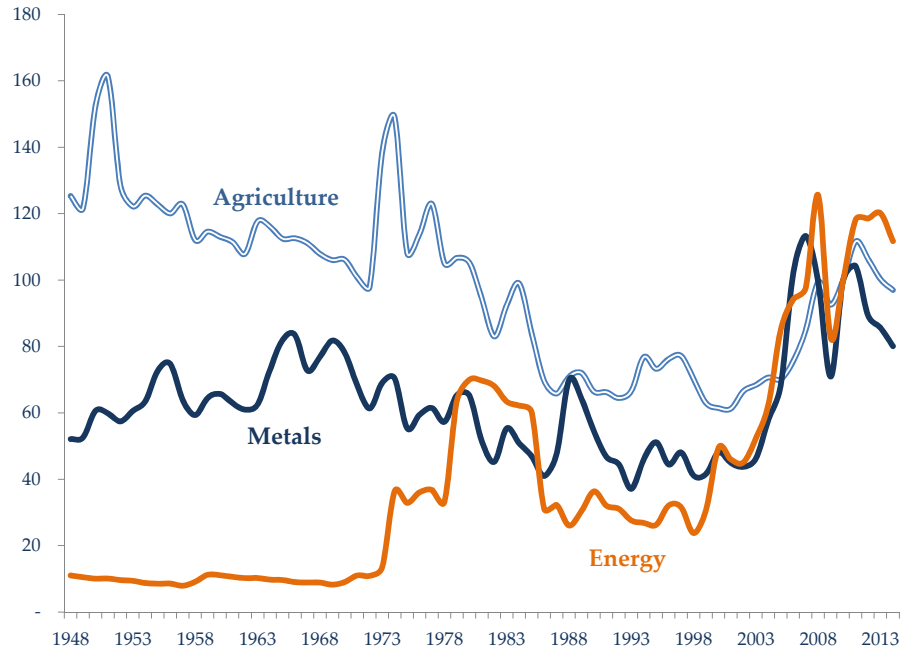
Notes: θ_1 is the long-run impact of income on nominal prices, θ_2 is the long-run impact of manufacturing prices on nominal prices, λ is the speed of adjustment to long-run equilibrium. The model is estimated based on equation (7) in text (the nominal price version). One (*), two (**), and three (***) asterisks indicate significance at 10%, 5%, and 1% levels, respectively.

Table 7
Short-Run Estimates from Panel Autoregressive Distributed Lag Models, per Capita Income, PPP terms

	---------- One lag ----------				---------- Two lags ----------			
	World	**High**	**Middle**	**Low**	**World**	**High**	**Middle**	**Low**
Terms of Trade (ToT)								
Constant	6.08***	6.00***	4.65***	3.39*	6.61***	6.75***	4.58***	2.99
	(7.29)	(6.46)	(8.01)	(1.89)	(6.71)	(7.70)	(5.37)	(1.53)
Stocks-to-use	-0.34***	-0.35***	-0.34***	-0.41***	-0.32***	-0.32***	-0.36***	-0.44***
	(-8.63)	(-7.29)	(-10.81)	(-7.29)	(-7.57)	(-6.71)	(-10.46)	(-7.08)
Ex. rate	-0.38***	-0.35***	-0.39***	-0.18	-0.41***	-0.37***	-0.37***	-0.16
	(-3.63)	(-3.40)	(-3.62)	(-0.99)	(-3.76)	(-3.40)	(-3.15)	(-0.89)
Int. rate	-0.01	-0.01	-0.00	-0.01	-0.01	-0.01*	-0.00	-0.01
	(-1.19)	(-1.63)	(-0.29)	(-1.14)	(-1.05)	(-1.93)	(-0.04)	(-0.94)
Oil Price	0.10***	0.11***	0.08**	0.02	0.11***	0.12***	0.08**	0.02
	(3.09)	(3.47)	(2.46)	(0.42)	(3.06)	(3.43)	(2.00)	(0.37)
GDP	3.36***	2.70***	2.25**	1.60***	3.94***	3.50***	2.17*	1.59***
	(3.59)	(4.35)	(2.53)	(2.94)	(4.36)	(6.04)	(1.83)	(2.61)
GDP (lag1)					0.55	-0.14	0.33	-0.33
					(0.82)	(-0.20)	(0.57)	(-0.56)
Nominal Prices								
Constant	5.30***	5.59***	4.77***	4.48***	5.33***	5.92***	4.97***	3.38***
	(6.88)	(6.28)	(12.52)	(5.06)	(5.26)	(6.51)	(7.68)	(2.91)
Stocks-to-use	-0.35***	-0.36***	-0.35***	-0.34***	-0.29***	-0.31***	-0.31***	-0.29***
	(-7.08)	(-6.34)	(-7.85)	(-6.14)	(-5.31)	(-5.67)	(-5.39)	(-4.39)
Ex. rate	-0.59***	-0.56***	-0.61***	-0.58***	-0.68***	-0.61***	-0.71***	-0.67***
	(-4.51)	(-3.64)	(-6.97)	(-5.40)	(-5.82)	(-4.98)	(-6.79)	(-5.77)
Int. rate	0.01	0.00	0.01***	0.01***	0.01***	0.01	0.02***	0.02***
	(1.55)	(0.70)	(3.44)	(2.96)	(2.60)	(1.14)	(4.78)	(3.95)
Oil Price	0.14***	0.14***	0.13***	0.14***	0.16***	0.16***	0.15***	0.16***
	(6.26)	(7.14)	(4.72)	(5.19)	(5.66)	(5.75)	(4.89)	(5.20)
GDP	1.52	1.67**	0.22	0.30	2.19***	2.58***	0.13	0.60
	(1.51)	(2.40)	(0.29)	(0.57)	(2.58)	(4.77)	(0.14)	(1.63)
GDP (lag1)					-0.74	-0.71	-0.55	-1.88***
					(-1.13)	(-1.12)	(-0.95)	(-3.08)
MUV	0.87***	0.80***	1.00***	1.01***	0.78***	0.68***	0.96***	0.91***
	(4.15)	(3.47)	(5.63)	(6.70)	(6.94)	(6.16)	(10.10)	(10.43)
MUV (lag1)					0.03	-0.06	-0.01	0.05
					(0.04)	(-0.91)	(-0.13)	(0.55)

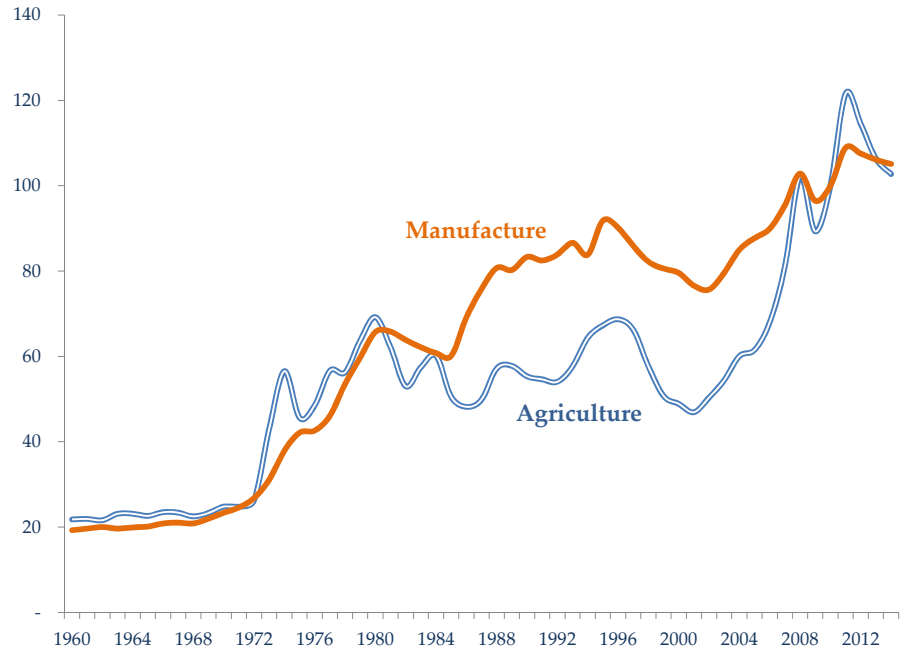
Notes: The parameter estimates are based on equation (7). The numbers in parentheses are *t*-statistics. One (*), two (**), and three (***) asterisks indicate significance at 10%, 5%, and 1% levels, respectively.

Figure 1
Ratio of Primary to Manufacturing Commodity Price Indices (2010 = 100)



Source: World Bank

Figure 2
Nominal Agricultural and Manufacturing Price Indices (2010 = 100)



Source: World Bank

Figure 3
A 2-Sector Model with Neutral Technical Change and Non-Homothetic Preferences

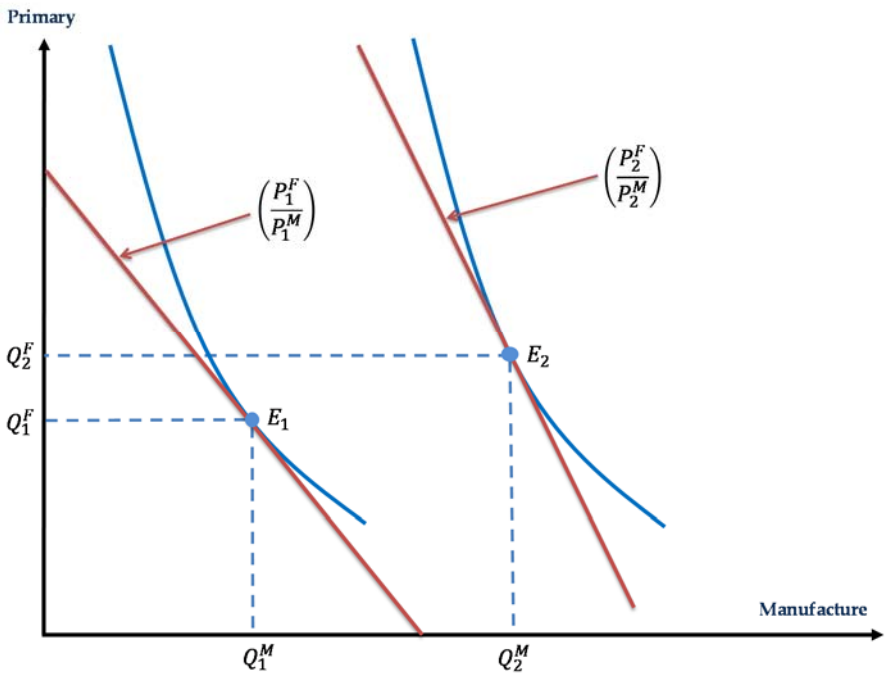
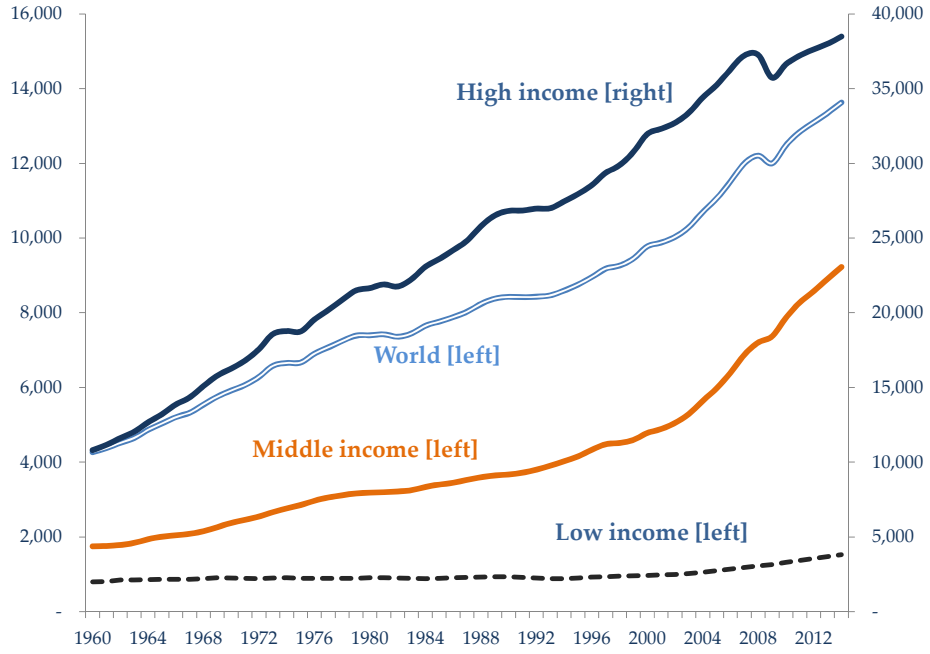


Figure 4
Per Capita Income, PPP Terms, Thousand $U.S.



Source: World Bank

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Appendix A: Summary of Empirical Research on ToT and the Prebisch-Singer Hypothesis

AUTHOR(S)	DATA	METHOD	MAIN CONCLUSION
Engel (1857)	Budget expenditures of 153 Belgian families, 1853	Applied non-parametric analysis	Poor families spend a larger proportion of their income on food than wealthier families.
Kindleberger (1943)	No data used	No model employed	ToT moves against primary commodity producing countries as income grows.
Prebisch (1950)	Primary commodity and manufacturing price indices, 1876-1947	Applied non-parametric analysis	The price ratio turned steadily against primary commodities from the 1870's until WWII.
Singer (1950)	1870s-WWII	No statistical analysis, made reference to the UN study	Prices turned against primary commodities and in favor of manufacturing goods.
Kindleberger (1958)	Unit value, country specific indices, 1872-1952	Applied non-parametric analysis	ToT moved against underdeveloped and in favor of developed countries.
Morgan (1959)	Primary & manufacturing price indices, 1860s to 1950s	Applied non-parametric analysis	Major fall (US), major rise (UK, New Zealand), mixed trend (India, Japan).
Spraos (1980)	Primary commodity prices, deflated by manufacturing price indices, 1871-1970	Tested for a linear trend	Exhibited a negative trend up to WWII
Sapsford (1985)	Primary commodity prices, deflated by manufacturing price indices, 1900-1982	Tested for a linear trend, allowing for structural breaks	Exhibited a negative trend after WWII subject to an upward shift in 1950.
Thirlwall and Bergevin (1985)	Primary commodity prices, deflated by manufacturing price indices, 1954-1982	Tested for a linear trend	Exhibited either constant or deteriorating trend.
Grilli and Yang (1988)	Index, sub-indices, and 24 prices, MUV-deflated, 1900–1983 (G-Y index)	Tested for a linear trend	All series exhibited negative trend.
Cuddington and Urzua (1989)	G-Y index, 1900-1983	Tested for a linear trend, allowing for structural breaks	Did not exhibit a negative trend except an abrupt drop after 1920.
von Hagen (1989)	G-Y index, 1900–1986	Tested for cointegration	Rejected the P-S hypothesis.
Helg (1991)	G-Y index, 1900-1988	Tested for a linear trend, allowing for structural breaks	Exhibited a negative trend after 1920.
Powell (1991)	G-Y index and sub-indices, 1900-1986	Tested for cointegration, accounting for jumps	Marginally in favor of stepwise declines in 1921, 1938, and 1975.
Sarkar and Singer (1991)	Country-specific ToT, 1965-1985	Tested for a linear trend	Exhibited a negative trend in some countries.
Boughton (1991)	Primary commodity prices, deflated by manufactured goods prices, 1854-1990	Tested for a linear trend though an error-correction model	Exhibited a negative trend.

(continued next page)

AUTHOR(S)	DATA	METHOD	MAIN CONCLUSION
Ardeni and Wright (1992)	G-Y index, 1900-1988	Applied trend-cycle decomposition	Exhibited a constant or declining trend.
Cuddington (1992)	G-Y individual prices plus oil & coal, 1900-1983	Tested for a linear trend, allowing for structural breaks	Most series did not exhibit a negative trend.
Sapsford, Sarkar, and Singer (1992)	Import and export prices of Britain & US, 1870-1938, and G-Y index, 1900-1986	Tested for a linear trend, allowing for structural breaks	All series exhibited a negative trend.
Bleaney and Greenaway (1993)	G-Y index and sub-indices, 1900-1991	Tested for a linear trend, allowing for structural breaks	Slow, long-run downward trend.
Lipsey (1994)	Various measures of manufacturing export prices, 1953-1991	Applied non-parametric analysis	The MUV overstates the long-run rise in manufactured goods prices.
Reinhart and Wickham (1994)	Quarterly indices, deflated by import unit values, 1959-1993	Used trend-cycle decomposition	Exhibited a negative trend that steepened in the 1980s.
Sapsford and Balasubramanyam (1994)	No data used	Reviewed statistical models	Broad support for the P-S hypothesis.
Newbold and Vougas (1996)	G-Y index, 1900-1992	Tested for a linear trend, allowing for structural breaks	No negative trend.
Bloch and Sapsford (1997)	Post-WWII price & wage data	Used structural model	Support for the P-S hypothesis.
Chen and Stocker (1997)	G-Y individual prices & real aggregate output, 1900-1986	Used structural model	Primary commodity prices not affected by a stabilization of Northern incomes.
Leon and Soto (1997)	G-Y index, sub-indices, & individual prices, 1900-1992	Tested for a linear trend, allowing for one structural break	Most series exhibited negative trends.
Lutz (1999)	G-Y index, 1900-1995	Tested for a linear trend	Exhibited a negative trend.
Bloch and Sapsford (2000)	Post-World War II price & wage data	Used structural model	Exhibited a negative trend.
Cashin and McDermott (2002)	Economist's commodity price index, 1862-1999	Tested for a linear trend	Exhibited a negative trend.
Blattman, Hwang, and Williamson (2003)	ToT & GDP data for 35 countries, 1870-1938	Applied panel regression	Secular ToT growth had a positive impact on growth performance.
Kim, Pfaffenzeller, Rayner, and Newbold (2003)	G-Y prices, 1900-1998	Tested for a linear trend	Some series exhibited negative trends.
Persson and Terasvirta (2003)	G-Y index, 1900-1995	Tested for trend with non-linear models	No negative trend.
Toye and Toye (2003)	No data used	Reviewed the literature	Examined origins & interpretation of P-S.

(continued next page)

AUTHOR(S)	DATA	METHOD	MAIN CONCLUSION
Bunzel and Vogelsang (2005)	G-Y index, 1900-1995	Tested for a linear trend	Exhibited a negative trend.
Newbold, Pfaffenzeller, and Rayner (2005)	G-Y index & individual prices, 1900-2002	Tested for a linear trend	Some series exhibited negative trends.
Zanias (2005)	G-Y index, 1900-1995	Tested for a linear trend, allowing for structural breaks	No negative trend, two structural breaks.
Kellard and Wohar (2006)	G-Y individual prices, 1900-1998	Tested for a linear trend, allowing for structural breaks	Some series exhibited negative trends for some sub-periods.
Svedberg and Tilton (2006)	Copper price, US CPI & PPI-deflated, 1870-2000	Tested for a linear trend, adjusting for inflation bias	No negative trend.
Molick, Faria, Albuquerque, and Leon-Ledesma (2008)	US PPI data, 1947-1999	Tested for a linear trend, allowing for structural breaks	Most series exhibited negative trends.
Balagtas and Holt (2009)	G-Y individual prices, 1900-2003	Applied forward simulation	Few series exhibited negative trend.
Cuddington (2010)	Copper price, US CPI-deflated, 1870-2006	Tested for a linear trend	No negative trend.
Harvey, Kellard, Madsen, and Wohar (2010)	25 commodity prices, MUV-deflated, 1650-2005	Tested for a linear trend, allowing for structural breaks	Some series exhibited negative trend.
Ghoshray (2011)	G-Y individual prices, 1900-2003	Tested for a linear trend, allowing for structural breaks	Most series exhibited negative trend for some sub-periods.
Fernandez (2012)	Various prices & indices, deflated by HPIM, MUV, PPI, and CPI, 1900-2008	Tested for a linear trend, allowing for structural breaks	Mixed evidence depending on time frequency & price deflator.
Arezki, Hadri, Kurozumi, and Rao (2012)	Nine commodity prices, US CPI-deflated, 1960-2007	Tested for a linear trend, allowing for structural breaks	Most series exhibited negative trend.
Erten and Ocampo (2013)	Five indices (extended G-Y data), & oil price, MUV-deflated, 1865-2010	Tested for super-cycles	A step-wise deterioration of real prices, supporting P–S hypothesis.
Yamada and Yoon (2014)	G-Y index, sub-indices, & individual prices, 1900-2010	Tested for piecewise linear trend	Most series exhibited negative trend for some sub-periods.
Arezki, Hadri, Loungani, Rao (2014)	Twenty five commodity prices, deflated by manufacturing prices, 1650-2010	Tested for piecewise linear trend, allowing for structural breaks	Most series exhibited negative trend.
Mariscal and Powell (2014)	G-Y index deflated by the MUV, 1900-2010	Tested for endogenous breaks	Little evidence of declining terms of trade, two structural breaks.

Notes: ToT = terms of trade, MUV = Manufacture Unit Value, CPI = Consumer Price Index, PPI = Producer Price Index, P-S = Prebisch-Singer. G-Y = Grilli-Young index, a trade-weighted average of 24 primary commodity prices at an annual frequency, deflated by MUV. In addition, Grilli and Yang (1988) constructed several sub-indices, including agricultural food commodity index, non-food agricultural commodity index, and metals. Subsequent studies that utilized the G-Y dataset have used the 24 individual commodity prices, various sub-indices, or the G-Y composite index.