# International Relative Price Levels: An Empirical Analysis

Charles Thomas, Jaime Marquez, Sean Fahle, and James Coonan

This chapter characterizes the structure of international relative price levels using cross-sectional and time series data.[1] The cross-sectional data consist of the 2005 benchmark purchasing power parities (PPPs) of the International Comparison Program (ICP) for 146 countries and 155 basic headings for products and services. These benchmarks embody methodological improvements enhancing product comparability and are the first to include actual price observations for China. The time series data consist of the PPPs reported by the 2008 World Development Indicators (WDI), which is the first set of indicators to use the new PPPs to revise the time series data. We use these data to measure bilateral relative price levels for 34 countries from 1980 to 2007. We then aggregate these bilateral relative prices with time-varying trade weights to obtain the evolution of a country's international relative price level.

In this chapter, we also use these data to answer questions of interest to international economists: Are the prices of a given product equalized across countries as one might expect based on the purchasing power parity hypothesis?[2] If not, is the resulting dispersion of relative prices across countries related to whether the product is internationally tradable? Furthermore, is there a systematic relation between the cross-country dispersion of relative price levels and that of income levels? If so, is that relation sensitive to whether the product is tradable? How comparable are our WDI-based measures of international relative prices with those based on the widely used Penn World Table (PWT), specifically Version 6.2? This comparison allows a quantitative assessment of the importance of the 2005 benchmarks.[3] Finally, how large are the differences between our measure of international relative prices and the widely used indexes of real effective exchange rates (REERs)? Such a comparison is relevant because existing REERs cannot, by design, capture the level effects from changes in the country composition of world trade.

Our analysis leads to several findings. First, the cross-country dispersion of prices at the level of basic headings depends importantly on whether the basic heading is tradable. Differences in the level of development across countries are relevant to explain the dispersion of prices. Previous research reached similar conclusions, but these findings confirm that the results hold up with more comparable and complete price data. Second, countries' aggregates of international relative prices based on the 2008 WDI data differ substantially from those based on the PWT 6.2 data. Finally, depictions of price movements based on our international relative prices are fundamentally different from those based on existing REERs.

# Cross-Sectional Evidence

## Data

The ICP provided the 2005 benchmark PPPs for 146 countries and 155 basic headings.[4] It also provided the 2005 values for gross domestic product (GDP), expenditures on each basic heading, population, and market exchange rates.[5]

The 2005 ICP benchmarks have two advantages over previous benchmarks. First, they are the first to include actual price observations for China and the first since 1985 to include actual price observations for India (the 1993 results for both countries were imputed). Indeed, as Deaton and Heston (2008) note, previous price data for these countries have been based on partial information and indirect methods. Second, the price collection for the 2005 benchmarks relied on the ICP's structured product descriptions, which is a list of standardized attributes used to identify a product as narrowly as possible (World Bank 2008, 142). This identification enhances the comparability of prices.[6]

## Empirical Results

Using this information, we measure the 2005 bilateral relative price level of the United States with respect to country $j$ in basic heading $i$ as

$$(22.1) \qquad q^{i}_{j,us} = \frac{E_{j/\$}}{PPP^{i}_{j}}; \qquad i = 1, \ldots, 92; j = 1, \ldots, 144$$

where $E_{j/\$}$ is the 2005 market exchange rate for country $j$ with respect to the U.S. dollar, and $PPP^{i}_{j}$ is the PPP exchange rate of the $i$-th basic heading in the $j$-th country. A value of two for $q^{i}_{j,us}$ means that the price level in the United States for the $i$-th basic heading is twice the price level of the same basic heading in country $j$.

Figure 22.1 shows the percentiles of the distribution of $\ln q^{i}_{j,us}$ across countries. To facilitate the presentation, we first split these distributions in two groups, tradable and nontradable, and then rank each group using its median.[7] The figure shows that the median of the distributions for nontradable products is generally higher than that for tradable products.[8] Also, the dispersion of relative prices for tradables is considerably smaller than the one for nontradables.

We now examine the extent to which the cross-country dispersion of relative prices is related to the cross-country dispersion of income levels.[9] To that end, we assume that

$$(22.2) \qquad \ln q^{i} = \alpha_{i} + \beta_{i} \cdot \ln y + u_{i}; \qquad i = 1, \ldots, 92; u_{i} \sim N(0, \sigma^{2}_{i})$$

where $q^{i} = q^{i}_{1,us}, \ldots, q^{i}_{j,us}, \ldots, q^{i}_{144,us}$; $y$ is a 144 × 1 vector of relative per capita GDPs, measured as $q^{i}_{144,us}$, the international dollar value of the $j$-th country's per capita GDP relative to that of the United States; and $u_{i}$ is a 144 × 1 vector of disturbances assumed to be white noise.

Figure 22.2 shows the estimates of $\beta_i$ and their 95 percent confidence bands. We arranged these estimates using the ordering of figure 22.1 as the template. The results indicate that these estimates are generally negative and significant. In other words, an *increase* in the per capita income of the *j*-th country relative to U.S. per capita income tends to lower the $q^i$ for that country, which corresponds to an *increase* in the price of the *i*-th good in the *j*-th country relative to the corresponding U.S. price. Note also that nontradables have the largest estimate of $\beta_i$ (in absolute value). This pattern has an economic explanation: an increase in income of the *j*-th country raises the demand for tradable and nontradable products, but the latter are supplied locally only. Hence an increase in the prices of these nontradable products drives them higher relative to U.S. prices. For tradable products, the existing forces to arbitrage prices are already reducing price differences with respect to the United States, and hence differences in the level of development are quantitatively less important.

Overall, the evidence from figures 22.1 and 22.2 suggests that differences in both the level of development and tradability are relevant considerations in explaining the cross-country dispersion of relative prices for the basic headings used here.

## Time Series Evidence

We now turn to the evolution of relative prices over time. To this end, we use the 2008 WDI data for purchasing power parities at the GDP level and market exchange rates for 1980–2007 for 29 countries.[10] An important advantage of these PPPs is that they rest on the ICP benchmarks, meaning that the comparisons across time and space rest on the same (high-quality) price data.

We begin by measuring the U.S. bilateral relative price with respect to the *k*-th country as

$$q_{us,k,t} = \frac{E_{k/\$,t}}{PPP_{k/us,t}} \tag{22.3}$$

where $E_{k/\$,t}$ is the market exchange rate between the *k*-th currency and the U.S. dollar, and $PPP_{k/us,t}$ is the corresponding PPP exchange rate reported by the WDI. Note that a value of two for $q_{us,k}$ means that the *basket* of products produced in country *k* is twice as expensive in the United States as in country *k*.

To measure relative prices for other countries, we exploit the transitivity of PPPs. Thus we estimate $q_{jk}$ as

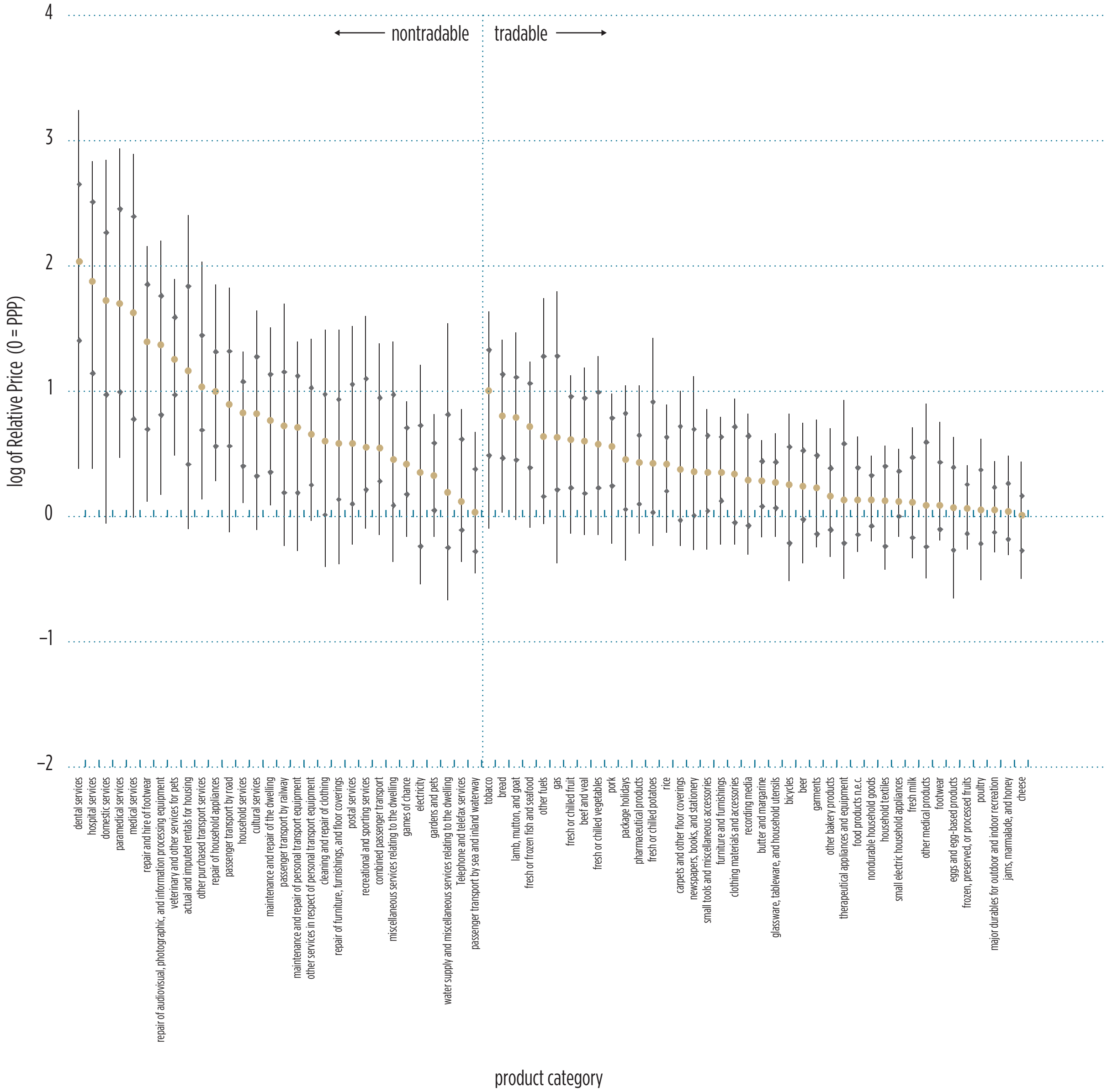
$$q_{jk} = \frac{q_{us,k}}{q_{us,j}} . \tag{22.4}$$

From these bilateral relative prices we obtain a multilateral measure using a geometric mean,

$$Q^g_{jt} = \prod_{\substack{k=1 \\ k \neq j}}^{29} (q_{jk,t})^{w_{jk,t}} , \tag{22.5}$$

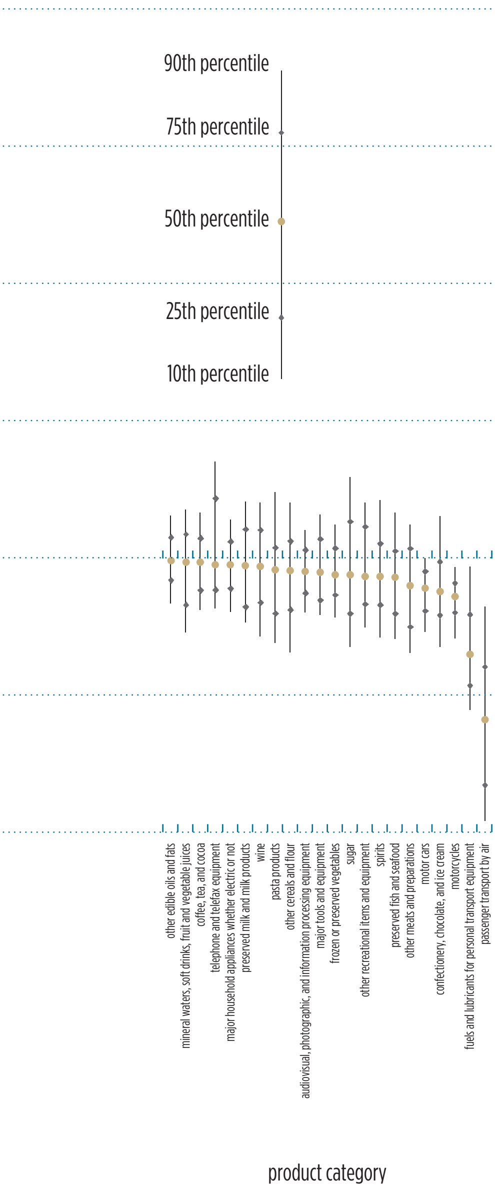
where $Q^g_{jt}$ is the international relative price of the *j*-th country, and $w_{jk,t}$ is the time-varying trade weight associated with the *k*-th country. For weights, we follow the method adopted by the

**FIGURE 22.1** Distribution of Relative Prices: Selected Basic Headings, 2005




Federal Reserve for its broad dollar index (Leahy 1998). These weights are designed to reflect the composition of world trade from the standpoint of the *j*-th country.[11] Equation (22.5) has two important properties. First, a value of two means that prices in the *j*-th country are twice as high as the average of its trading partners. Second, even if prices levels were fixed, the $Q^g$ for each country changes as the composition of world trade changes.[12]

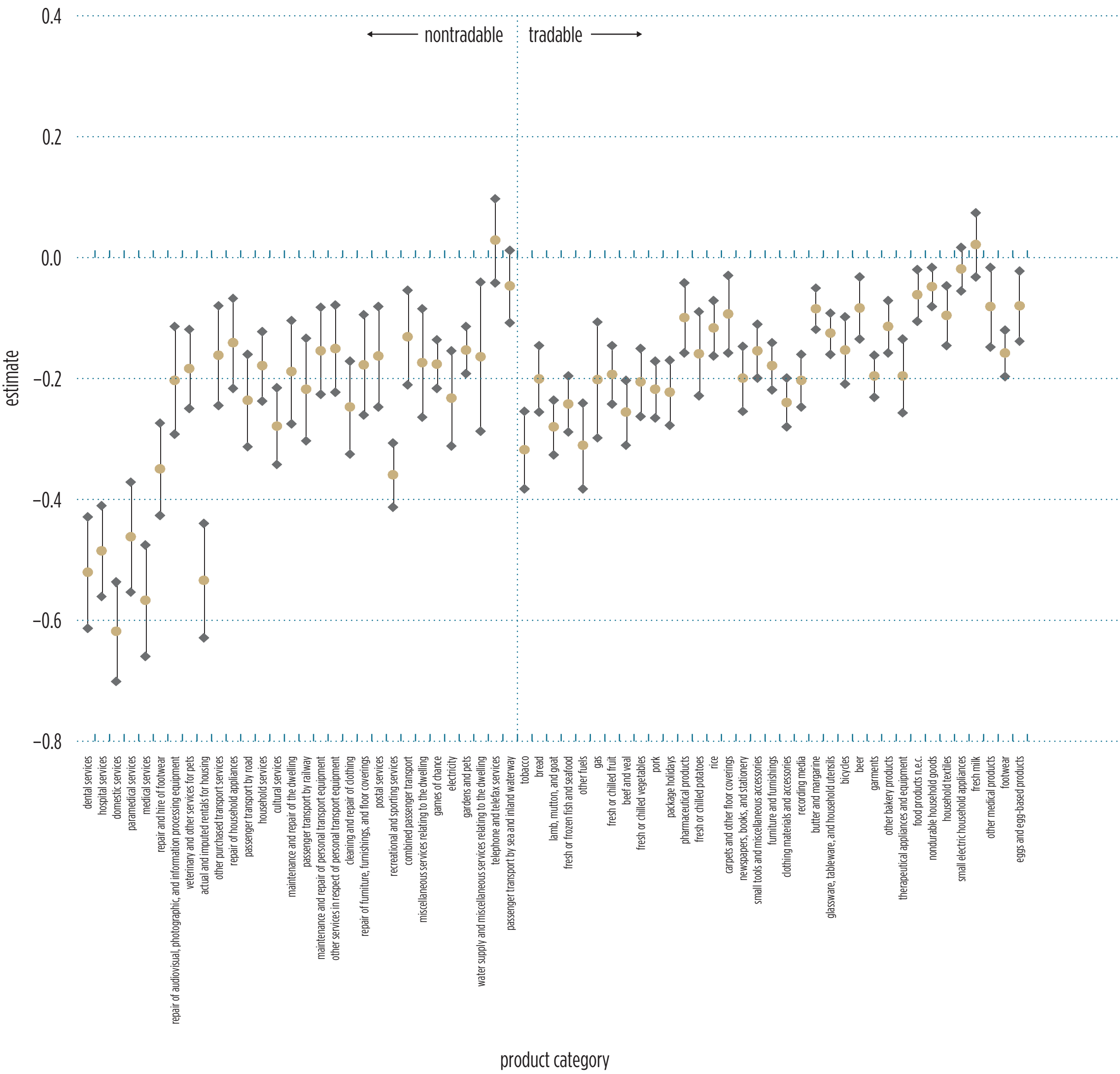
The next section discusses how this measure of relative prices differs across countries. It is followed by a section that compares the evolution of this measure to the more familiar chained indexes of real effective exchange rates.




## International Relative Prices

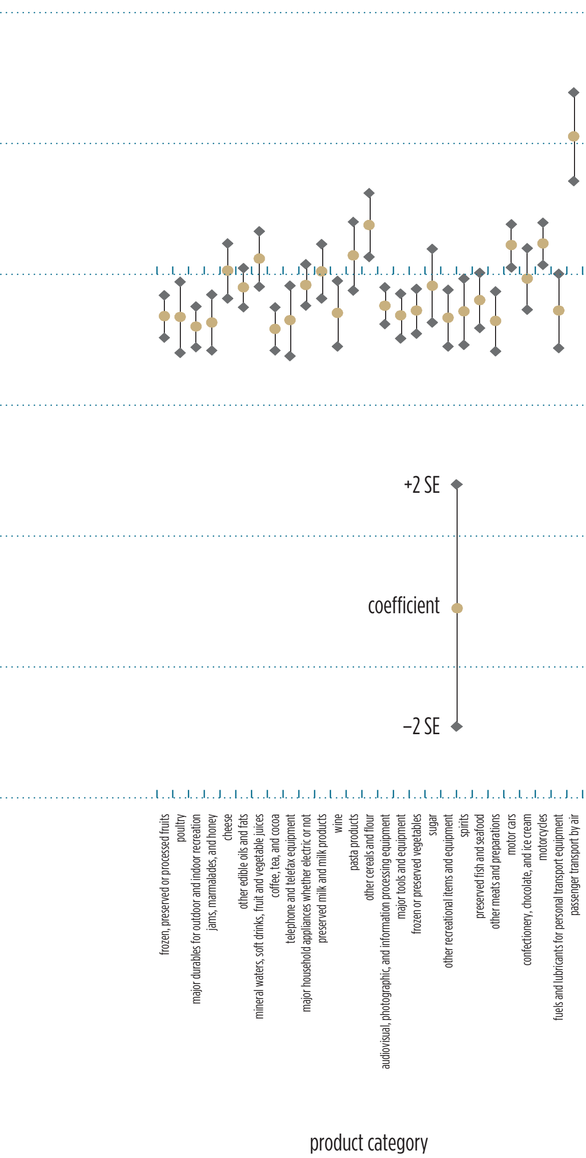
Our calculations for $Q^g$ (table 22.1) show that the WDI-based measures of international relative prices differ markedly across countries. For example, the prices in 2007 (column [6]) for China and India are less than half those of their trading partners, whereas Japan's prices are about 50 percent higher than those of its trading partners. The U.S. international relative prices, which are among the highest, were the highest during the mid-1980s. Moreover, price measures based on the 2008 WDI differ significantly from those based on PWT 6.2. For example, the 2007 international

**FIGURE 22.2** Confidence Intervals (95 Percent Level) for Estimate of $\beta_i$ in Equation (22.2)




relative price for China is 0.50 using WDI data and 0.33 using PWT data, a gap of 52 percent. Thirteen countries have an average price revision of at least 20 percent (column [13]). The revisions are also noticeable for countries with significant trade with China: 18 percent for Japan and 13 percent for Korea. A few countries have relatively small revisions: Mexico, the United Kingdom, and the United States.

Finally, international relative prices have increased since 1995 for most countries. Many factors are responsible for this pattern, but two of them are of interest. First, the development of emerging market countries, with the accompanying increase in the demand for nontradables, raises those countries' general price levels, everything else unchanged. Second, the rapid expansion of




exports of low-priced products from emerging market economies lowers the average world price of tradables, and, because these exports are a large share of industrial countries' imports, raises their prices relative to world prices.

## Real Effective Exchange Rate Indexes

We now compare our WDI-based international relative prices with existing measures of real effective exchange rate indexes. This comparison is of interest because most analyses of the role of international competitiveness in external imbalances use chained indexes of real effective exchange rates.

**TABLE 22.1** International Relative Prices for WDI and PWT, Quarter 1 of Selected Years

| Country | WDI | | | | | | PWT | | | | | | Average revision, 1995–2007[b] |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | 1980q1 (1) | 1985q1 (2) | 1990q1 (3) | 1995q1 (4) | 2000q1 (5) | 2007q1 (6) | 1980q1 (7) | 1985q1 (8) | 1990q1 (9) | 1995q1 (10) | 2000q1 (11) | 2007q1[a] (12) | (13) |
| Argentina | 0.56 | 0.80 | 0.58 | 1.11 | 1.12 | 0.62 | 1.53 | 0.76 | 0.49 | 0.86 | 0.94 | *0.51* | 23.3 |
| Australia | 1.18 | 1.30 | 1.05 | 1.01 | 1.08 | 1.37 | 1.26 | 1.30 | 1.18 | 1.07 | 1.12 | *1.57* | −7.3 |
| Austria | 0.92 | 0.90 | 0.98 | 1.05 | 1.01 | 1.07 | 0.95 | 0.83 | 1.05 | 1.32 | 1.10 | *1.43* | −17.9 |
| Belgium | 1.17 | 0.87 | 0.96 | 1.06 | 1.04 | 1.13 | 1.33 | 0.91 | 1.17 | 1.41 | 1.17 | *1.55* | −21.0 |
| Brazil | 0.51 | 0.53 | 0.69 | 0.62 | 0.71 | 0.72 | 0.43 | 0.44 | 0.69 | 0.71 | 0.58 | *0.78* | 0.7 |
| Canada | 1.07 | 1.08 | 1.10 | 0.95 | 0.90 | 1.15 | 1.03 | 1.16 | 1.17 | 0.97 | 0.99 | *1.25* | −6.4 |
| Chile | 0.92 | 0.77 | 0.52 | 0.61 | 0.69 | 0.81 | 0.70 | 0.59 | 0.47 | 0.54 | 0.58 | *0.64* | 19.5 |
| China | 0.84 | 0.74 | 0.40 | 0.35 | 0.45 | 0.50 | 0.70 | 0.45 | 0.27 | 0.22 | 0.27 | *0.33* | 59.1 |
| Finland | 1.03 | 1.13 | 1.35 | 1.12 | 1.23 | 1.36 | 1.04 | 1.12 | 1.41 | 1.58 | 1.46 | *1.76* | −22.5 |
| France | 1.21 | 1.06 | 1.15 | 1.17 | 1.12 | 1.16 | 1.31 | 0.99 | 1.23 | 1.34 | 1.12 | *1.42* | −10.3 |
| Germany | 1.17 | 0.97 | 1.09 | 1.23 | 1.18 | 1.18 | 1.27 | 0.93 | 1.21 | 1.48 | 1.22 | *1.54* | −14.5 |
| India | 0.62 | 0.64 | 0.46 | 0.33 | 0.38 | 0.41 | 0.39 | 0.39 | 0.26 | 0.21 | 0.23 | *0.27* | 58.1 |
| Indonesia | 0.79 | 0.77 | 0.44 | 0.43 | 0.33 | 0.60 | 0.43 | 0.42 | 0.27 | 0.27 | 0.24 | *0.38* | 51.6 |
| Ireland | 0.90 | 0.92 | 0.98 | 0.98 | 1.05 | 1.23 | 1.02 | 0.98 | 1.13 | 1.24 | 1.16 | *1.59* | −17.7 |
| Italy | 0.74 | 0.84 | 0.98 | 0.90 | 0.97 | 1.15 | 0.92 | 0.84 | 1.18 | 1.09 | 1.14 | *1.58* | −19.9 |
| Japan | 1.19 | 1.12 | 1.70 | 2.11 | 1.86 | 1.55 | 1.17 | 1.29 | 1.65 | 2.54 | 2.45 | *1.79* | −18.1 |
| Korea, Rep. | 0.69 | 0.67 | 0.71 | 0.82 | 0.71 | 1.04 | 0.69 | 0.69 | 0.72 | 0.84 | 0.88 | *1.24* | −12.6 |
| Malaysia | 0.86 | 0.84 | 0.58 | 0.56 | 0.50 | 0.65 | 0.71 | 0.74 | 0.52 | 0.51 | 0.45 | *0.58* | 11.0 |
| Mexico | 0.63 | 0.56 | 0.46 | 0.50 | 0.61 | 0.75 | 0.71 | 0.66 | 0.54 | 0.44 | 0.77 | *0.91* | −8.2 |

| | | | | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Netherlands | 1.23 | 1.04 | 1.06 | 1.09 | 1.04 | 1.22 | 1.33 | 0.95 | 1.18 | 1.38 | 1.21 | *1.76* | −21.9 |
| Philippines | 0.56 | 0.54 | 0.46 | 0.50 | 0.51 | 0.54 | 0.38 | 0.38 | 0.31 | 0.34 | 0.33 | *0.38* | 47.9 |
| Portugal | 0.55 | 0.53 | 0.60 | 0.77 | 0.80 | 0.89 | 0.57 | 0.47 | 0.61 | 0.76 | 0.71 | *1.05* | −0.4 |
| Singapore | 0.75 | 0.86 | 0.74 | 0.94 | 0.97 | 0.93 | 0.82 | 1.00 | 0.91 | 1.13 | 1.12 | *1.24* | −18.4 |
| Spain | 0.86 | 0.70 | 0.85 | 0.85 | 0.83 | 0.93 | 1.00 | 0.73 | 1.04 | 1.04 | 0.94 | *1.38* | −20.9 |
| Sweden | 1.38 | 1.16 | 1.31 | 1.18 | 1.22 | 1.25 | 1.40 | 1.22 | 1.39 | 1.23 | 1.32 | *1.41* | −7.7 |
| Switzerland | 1.27 | 1.18 | 1.27 | 1.40 | 1.35 | 1.40 | 1.27 | 1.20 | 1.36 | 1.62 | 1.45 | *1.53* | −9.7 |
| Thailand | 0.68 | 0.67 | 0.55 | 0.57 | 0.51 | 0.55 | 0.46 | 0.45 | 0.39 | 0.40 | 0.39 | *0.51* | 27.0 |
| United Kingdom | 0.85 | 0.88 | 0.96 | 0.97 | 1.17 | 1.31 | 1.00 | 0.83 | 0.97 | 0.96 | 1.24 | *1.38* | −3.2 |
| United States | 1.02 | 1.48 | 1.08 | 1.10 | 1.31 | 1.31 | 1.06 | 1.48 | 1.09 | 1.12 | 1.31 | *1.37* | −2.1 |

*Source:* Data from World Development Indicators and Penn World Table.
*Note:* WDI = World Development Indicators; PWT = Penn World Table.
a. Based on extrapolations of PWT 6.2 through 2007.
b. For the years shown in the table.

However, as shown shortly, these chained indexes cannot capture the direct effects that changes in the country composition of world trade have on international relative price levels.

To see this, note that the most widely used alternative to $Q^g_{j,t}$ is the chained aggregate, which is a weighted average of the growth rates of bilateral relative prices:

$$\frac{Q^c_{jt}}{Q^c_{j,t-1}} = \prod_{\substack{k=1 \\ k \neq j}}^{34} \left( \frac{\frac{P_{k,t}}{P_{k,t-1}} \cdot \frac{E_{j/k,t}}{E_{j/k,t-1}}}{\frac{P_{j,t}}{P_{j,t-1}}} \right)^{w_{jk,t}} \tag{22.6}$$

where $P_{k,t}$ is an "arbitrary" measure of prices in the $k$-th country, $P_{j,t}$ is an "arbitrary" measure of prices in the $j$-th country, and $E_{j/k,t}$ is the price of the currency of the $k$-th country in terms of the currency of the $j$-th country. This formulation is appealing because it relies on growth rates without having to measure price levels as such. For example, the term $\frac{P_{k,t}}{P_{k,t0}}$ could be the consumer price index (CPI) for the $k$-th country with $t_0$ as the base period. Indeed, the Bank for International Settlements (BIS) REER, which we use as a comparator, is based on CPIs (see Klau and Fung 2006). We use the BIS measure, which we denote as $Q^{bis}_{j,t}$, because it is well known and available for many emerging market economies.[13]

To facilitate our comparison with the BIS REER, we note that price series analogous to the CPIs, but based on GDP baskets, are implicit in the $q_{jk,t}$'s used in constructing $Q^g_{j,t}$. A suitable rearrangement of terms in equation (22.6) yields

$$\frac{Q^c_{jt}}{Q^c_{j,t-1}} = \prod_{\substack{k=1 \\ k \neq j}}^{29} \left( \frac{\frac{P_{k,t}}{P_{j,t}} \cdot E_{j/k,t}}{\frac{P_{k,t-1}}{P_{j,t-1}} \cdot E_{j/k,t-1}} \right)^{w_{jk,t}} = \prod_{\substack{k=1 \\ k \neq j}}^{29} \left( \frac{q_{jk,t}}{q_{jk,t-1}} \right)^{w_{jk,t}}. \tag{22.7}$$

Because this $Q^c_{j,t}$ is based on the same underlying price data as $Q^g_{j,t}$, we can more easily identify those differences attributable to the aggregation method as opposed to those attributable to the use of different underlying price data.

Specifically, logarithmic differentiation of $Q^c_{j,t}$ and $Q^g_{j,t}$ yields

$$\hat{Q}^g_{jt} = \sum_k w_{jk,t} \cdot dlnq_{jk,t} + \sum_k dw_{jk,t} \cdot lnq_{jk,t} \tag{22.8}$$

and

$$\hat{Q}^c_{jt} = \sum_k w_{jk,t} \cdot dlnq_{jk,t} \tag{22.9}$$

where ˆ denotes a growth rate. Thus if prices are constant, then $\hat{Q}^c_{jt} = 0$ necessarily, whereas $\hat{Q}^g_{jt}$ could differ from zero. Furthermore, the difference in growth rates between the geometric and the chained aggregate is

$$\hat{Q}^g_{jt} - \hat{Q}^c_{jt} = \sum_k dw_{jk,t} \cdot lnq_{jk,t}, \tag{22.10}$$

that is, if the weights are constant, then the two growth rates are identical. But if the weights are not constant, then the difference in growth rates reflects the interaction between each period's distribution of the level of bilateral relative prices and the evolution of the weights.

Table 22.2 compares the cumulative growth rates of $Q_t^g$ and $Q_t^{bis}$ from 1995 to 2007 (note that the subscript *j* has been dropped). Our calculations indicate that China's international relative price increases by 42 percent if one uses $Q_t^g$ (column [1]) and 18 percent if one uses $Q_t^{bis}$ (column [2]).

**TABLE 22.2** Growth Rates of International Relative Prices, 1995–2007

| | Growth rates | | | Growth rate differentials | | |
|---|---|---|---|---|---|---|
| Country | $Q^g$ (1) | $Q^{bis}$ (2) | $Q^c$ (3) | $Q^g - Q^{bis}$ (4) | $Q^g - Q^c$ (5) | $Q^c - Q^{bis}$ (6) |
| Argentina | −44 | −49 | −52 | 5 | 8 | −3 |
| Australia | 36 | 21 | 16 | 15 | 20 | −5 |
| Austria | 2 | −6 | −4 | 8 | 6 | 2 |
| Belgium | 7 | −5 | 0 | 11 | 6 | 5 |
| Brazil | 15 | −14 | 3 | 30 | 13 | 17 |
| Canada | 21 | 16 | 11 | 5 | 10 | −5 |
| Chile | 33 | −2 | 20 | 35 | 13 | 22 |
| China | 42 | 18 | 32 | 24 | 10 | 14 |
| Finland | 21 | −9 | 8 | 31 | 13 | 17 |
| France | −1 | −5 | −7 | 4 | 6 | −2 |
| Germany | −4 | −12 | −12 | 7 | 8 | 0 |
| India | 24 | 3 | 8 | 21 | 17 | 5 |
| Indonesia | 39 | −8 | 15 | 47 | 24 | 23 |
| Ireland | 26 | 20 | 23 | 6 | 3 | 3 |
| Italy | 28 | 15 | 16 | 13 | 12 | 1 |
| Japan | −26 | −37 | −35 | 11 | 8 | 3 |
| Korea, Rep. | 27 | 9 | 6 | 18 | 21 | −3 |
| Malaysia | 17 | −13 | 0 | 30 | 17 | 13 |
| Mexico | 50 | 63 | 35 | −13 | 15 | −28 |
| Netherlands | 12 | 0 | −1 | 11 | 13 | −1 |
| Philippines | 7 | −8 | −7 | 15 | 14 | 1 |
| Portugal | 16 | 7 | 8 | 9 | 8 | 1 |
| Singapore | −1 | −9 | −15 | 8 | 14 | −5 |
| Spain | 9 | 12 | 3 | −3 | 6 | −9 |
| Sweden | 6 | −9 | −1 | 15 | 7 | 7 |
| Switzerland | 0 | −15 | −7 | 15 | 7 | 8 |
| Thailand | −3 | −6 | −18 | 4 | 15 | −11 |
| United Kingdom | 35 | 20 | 27 | 16 | 9 | 7 |
| United States | 19 | 7 | 3 | 12 | 16 | −4 |

*Note:* Calculations based on the dates used in table 22.1.

Seven additional countries have gaps of at least 20 percent (column [4]). Furthermore, $Q^g$ and $Q^{bis}$ move in opposite directions, in our sample, for 10 countries: the differences in measures are not solely about magnitudes. These two findings suggest, then, that our characterization of international relative prices is fundamentally different from the one implied by the BIS measure.

Several reasons could be given for the difference in growth rates between $Q_t^g$ and $Q_t^{bis}$: aggregation methods, price measures, and country weights. A simple decomposition shows that most of the difference in growth rates can be attributed to differences in aggregation methods. In doing so, we express the gap in growth rates between $Q_t^g$ and $Q_t^{bis}$ as

$$\hat{Q}_t^g - \hat{Q}_t^{bis} = \left(\hat{Q}_t^g - \hat{Q}_t^c\right) + \left(\hat{Q}_t^c - \hat{Q}_t^{bis}\right). \quad (22.11)$$

The $Q_t^c$ used here is based on the same relative prices and weights as $\hat{Q}_t^g$. The only difference between these two measures can be attributed to the choice of aggregation method—that is, the difference between a geometric aggregate and a chain-weighted aggregate. Thus $(\hat{Q}_t^g - \hat{Q}_t^c)$ captures the effects of aggregation methods alone. The term $(\hat{Q}^c - \hat{Q}^{bis})$ captures the importance of the remaining factors: price measures and weighing schemes. Column (5) in table 22.2 shows that the difference between $\hat{Q}^g$ and $\hat{Q}^c$ explains most of the gap between $\hat{Q}^g$ and $\hat{Q}^{bis}$.[14] In other words, we find that the interactions between price levels and the structure of trade are sufficiently important to induce divergences between $Q_t^{bis}$ and $Q_t^g$.

## Conclusion

This chapter characterizes the distributions of relative price levels across countries, products, and time. We begin by studying the cross-country distributions of 2005 relative price levels across 144 countries for 92 detailed product categories. We find that the cross-country dispersion of relative price levels depends importantly on whether the basic heading is tradable. Differences in the level of development are also relevant for explaining the dispersion of prices.

We continue our analysis with a study of time series PPP data from the World Development Indicators and exchange rate data for 34 countries spanning 1980–2007. The WDI data have the benefit of being derived from the ICP benchmarks. We use these components to construct bilateral relative price levels, which are then aggregated using a geometric mean and weighted using time-varying trade weights. We find that countries' aggregates of international relative prices based on the 2008 WDI data differ substantially from those based on data from Version 6.2 of the Penn World Table. Finally, depictions of price movements based on our international relative prices are fundamentally different from those based on existing REERs. We do not interpret these divergences as a call to abandon existing effective exchange rate indexes. Rather, we interpret these divergences as an opportunity for $Q_t^g$ to complement the information in those indexes, a role that is likely to be present so long as changes in the pattern of trade continue. In this case, the ongoing efforts by the International Comparison Program are central to understanding international relative prices.

NOTES

1. The views in this chapter are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System or of any other person associated with the Federal Reserve System. Preliminary versions of this

paper were presented at George Washington University, the meetings of the fall 2009 Midwest International Economics Group (Penn State), and the fall 2009 Workshop of the Federal Reserve Board. We are grateful to Nada Hamadeh for providing the ICP data and to both Frederic A. Vogel and D. S. Prasada Rao for numerous comments on an earlier draft of this chapter.

2. This hypothesis indicates that, in the absence of transportation costs and government regulations, arbitrage equalizes the price for a given product across all countries when expressed in a common currency. For reviews, see Froot and Rogoff (1995), Taylor (2003), Chinn (2005), and Klau and Fung (2006). Other relevant papers include those by Lipsey, Molinari, and Kravis (1990); Hooper and Richardson (1991); and Turner and Van't dack (1993).
3. For details on the PWT, see Heston, Summers, and Aten (2006). For an introduction, see Summers and Heston (1991) and Gulde and Schulze-Ghattas (1993). Our comparison focuses on PWT 6.2 for two reasons. First, it allows us to compare results described in this chapter to previous work such as that by Thomas, Marquez, and Fahle (2008). Second, PWT 6.3 does not incorporate price data for China, and so little is lost by focusing on PWT 6.2.
4. A *basic heading* is the lowest level of disaggregation for which PPPs are computed (World Bank 2008, 14). An example of a basic heading is "confectionery, chocolate, and ice cream." These PPPs are constructed to equalize the dollar price of the associated basic heading across countries. Chapters 6 and 7 of the World Bank's *Global Purchasing Power Parities and Real Expenditures: 2005 International Comparison Program* describe the methodology used in the computation of PPPs at the basic heading level (World Bank 2008).
5. The data file received had incomplete data for Zambia and Zimbabwe, and so they are excluded from our analysis. Also, in response to suggestions from ICP staff, we excluded those items associated with government activities such as government production of health services, collective services, and social protection, because the cross-country comparability of these items is not sufficient for the purposes of this chapter.
6. See Chen and Ravallion (2008), appendix G of World Bank (2008), and Deaton and Heston (2008).
7. The tradable and nontradable split is, admittedly, ad hoc and based on the authors' *a priori* views. Thus further work is needed to assess the robustness of these results.
8. Tobacco and gas are, however, important exceptions because they are tradable products but have the largest relative prices—indeed, higher than many of the relative prices for nontradable products. One possible explanation for this seemingly odd result is that tobacco and gas are taxed at lower rates in the United States than in other countries.
9. See Obstfeld and Rogoff (1996) for a review.
10. Specifically, we use the 29 countries (or economies) included in the broad measure of the Federal Reserve's real effective value of the dollar (Leahy 1998): Argentina, Australia, Austria, Belgium, Brazil, Canada, Chile, China, Finland, France, Germany, India, Indonesia, Ireland, Italy, Japan, Republic of Korea, Malaysia, Mexico, the Netherlands, the Philippines, Portugal, Singapore, Spain, Sweden, Switzerland, Thailand, United Kingdom, and United States.
11. Specifically, the un-normalized broad weight for a given country is $\omega_{jk,t} = 0.5 \cdot \mu_{jk,t} + 0.25 \cdot \xi_{jk,t} + 0.25 \cdot \varsigma_{jk,t}$, where $\mu_{jk,t}$ is the share of imports from the $k$-th country, $\xi_{jk,t}$ is the export share to the $k$-th country, and $\varsigma_{jk,t}$ is the extent to which exports to the $k$-th country compete with exports from other countries. The normalized broad weight of the $j$-th country relative to the $k$-th country is $w_{jk,t} = \frac{\omega_{jk,t}}{\sum_i \omega_{jk,t}}$. We use data from the International Monetary Fund's Direction of Trade Statistics.

12. The exception is the case in which all relative prices are fixed at the value of one.
13. $Q_t^{bis}$ uses time-varying weights reflecting world trade shares in manufactures.
14. In 16 countries, the gap between $\hat{Q}^g$ and $\hat{Q}^c$ is more than twice the absolute value of the gap between $\hat{Q}^{bis}$ and $\hat{Q}^c$.

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