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**THE *REAL* REAL PRICE OF NONRENEWABLE
RESOURCES: COPPER 1870-2000**

by

Peter Svedberg and John E. Tilton

INSTITUTE FOR INTERNATIONAL ECONOMIC STUDIES
Stockholm University

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ABSTRACT

Over the past 40 years economists have devoted considerable effort to estimating long-run trends in real commodity prices. The results indicate that the real prices for many commodities have fallen, suggesting to the surprise of many that resource scarcity is declining over time. Almost all of this work, however, uses the U.S. producer price index or other standard price deflators, which recent research shows overestimate inflation for several reasons. This article examines copper prices with adjusted deflators designed to eliminate this bias. It finds that the trend over time, which is significantly downward when no adjustment is made to the deflator, displays no tendency in either direction or is significantly upward depending on the magnitude of the deflator adjustment employed. These findings suggest that real resource prices provide less support than widely assumed for the hypothesis that resources are becoming more available or less scarce over time.

Keywords: resources, copper, real price, inflation bias, scarcity

JEL classification: C3, F1, O13, Q3

* Peter Svedberg is Professor of Development Economics at the Institute for International Economic Studies, Stockholm University (peter.svedberg@iies.su.se). John E. Tilton is the William Jesse Coulter Professor of Mineral Economics, Division of Economics and Business, Colorado School of Mines (jtilton@mines.edu), and Profesor de la Catedra de Economia de Minerales, Centro de Minería, Pontificia Universidad Católica de Chile. They are grateful to Phillip Crowson, Peter Howie, and Marian Radetzki for helpful comments on an earlier draft, and to Alejandro D. Lombardia and John Tyler Hodge for their research assistance. Peter Svedberg acknowledges financial support from the Swedish International Development Cooperation Agency (*Sida*).

1. INTRODUCTION

Copper, oil, and many other natural resources are nonrenewable at least on any time horizon of relevance to the human race. As the easiest-to-find and richest deposits are exploited, many anticipate that the real price of mineral commodities will rise reflecting their increasing scarcity. However, dozens of empirical investigations using modern econometric techniques have failed to detect statistically significant upward trends in the long-run real price of mineral commodities, as Krautkremer (1998) notes in his review of this literature. The standard explanation is that new technologies, the discovery of new deposits, recycling, and other forces have more than offset the cost-increasing effects of depletion over the past 130 years.

In this earlier work, nominal prices are almost always reported in U.S. currency, and converted to real prices using the U.S. producer price index and other standard deflators. The choice of deflator is rarely explained in any detail, and in a few cases not even identified. Krautkremer (1998), for example, in his extensive review of the literature only mentions his choice of deflator (the U.S. consumer price index) in a footnote. A more serious problem, which provides the motivation for this study, is that a growing number of studies (see Table 1) have found that the traditional price deflators overstate inflation. This paper reassesses the long-run real price trends for one nonrenewable resource, copper, using deflators designed to correct for this bias.

Copper is a good choice for a pilot study for several reasons. First, it is an important metal. Recently the world has produced 13-15 million tons of refined copper annually (World Bureau of Metal Statistics, annual), worth nearly 25 billion dollars. Only the iron and steel, aluminium and gold industries have higher sales. Second, there are reliable and consistent data on annual copper prices back to 1870 (and even before). Third, copper is a relatively homogeneous good. Unlike petroleum, for example, with its many grades, refined copper is 99.99+ per cent pure copper.¹ Fourth, the cost of

¹ The London Metal Exchange copper price, for example, was for Trough Copper (fire-refined averaging 99.25 per cent copper) from 1780 to 1879, for Best Selected Copper (fire-refined averaging 99.75 per cent copper) from 1880 to 1914, and for electrolytic wirebars after 1914 (averaging 99.99 per cent copper). In the absence of these quality improvements, the price of copper presumably would have fallen slightly more or risen slightly less. However, some changes in copper trading have had the opposite effect. For instances, for technical reasons, sellers since 1981 have been able to deliver cathode copper instead of wirebar. This has eliminated a step in the production process and reduced costs. The U.S. producer price for copper reflects similar developments over time.

transporting copper is a small percentage of its final price. Fifth, copper is sold on global markets rather than in various regional markets. As a result, the prices in the two main markets, the US and London Metal Exchange (LME), move together within the bounds set by the cost of transporting copper across the Atlantic. Over the entire 1870-2000 period, the simple correlation coefficient between the LME and U.S. producer price, both expressed in U.S. dollars and deflated by the conventional CPI, is 0.90. Sixth, the copper market is reasonably competitive. During certain periods, collusive activities by producers or price controls by governments have affected prices. But, thanks to a number of scholars—Herfindahl (1959) and Howie (2002), for example—who have studied the copper industry, these periods are well documented, allowing one to control for their influence.²

The analysis begins in Section 2 with a brief review of the theory and previous empirical evidence on long-run trends in resources prices. Section 3 presents the reasons why the standard deflators overestimate inflation during recent decades, and hence the adjustments required to eliminate the bias. Section 4 considers the inflation bias before 1950. Section 5 estimates the *real*, real price of copper and carries out a series of robustness tests. Section 6 highlights the main findings, examines some of the implications, raises a few caveats, and makes suggestions for further research.

2. RESOURCE PRICE CHANGES: THEORY AND EVIDENCE

One of the main reasons for assessing long-term price trends for nonrenewable commodities is the concern that depletion over time is increasing their scarcity. Real prices are the most readily available of the few useful indicators of trends in resource availability.³

The formal theory in this area dates back to Hotelling (1931). His seminal article considers the optimal output over time for a mine with a fixed quantity of mineral reserves. Hotelling shows that the mine will behave differently than other firms whose production is not ultimately limited by an exhaustible resource. In particular, it will

² Howie (2002, ch.4) corrects for price distortions introduced by subsidization and the exercise of market power over the 1900-1997 period, and finds this produces less of a decline in the long-run trend in real copper prices.

expand its output during any period only up to the point where the cost of producing an additional unit plus user costs equals the market price, where user costs reflect the present value of the future profits forgone by producing an additional unit of output now.

Hotelling also demonstrates that user costs must rise by r per cent per year, where r per cent is the market rate of return on investments that are of similar risk and in other ways comparable to an investment in mineral reserves. If this were not the case, if for example the return on comparable investments were higher, then mine owners would have an incentive to exploit all of their reserves during the current period (Hotelling imposes no constraints on output other than the total quantity of reserves) and invest the profits in other investments. This behavior would raise the price of the mineral commodity in the future and so increase user costs. Not until the growth in user costs was comparable to the return on other similar assets, would mines have an incentive to stop diverting output from the future to the present. Should the pendulum swing too far, so that the user costs increased at more than r per cent a year, mines would want to hold onto their reserves. The diversion of output from the present to the future would slow the pace at which user costs rise until the latter was growing at r per cent a year.

Some students of Hotelling argue that his analysis implies that the prices of mineral commodities will rise at r per cent a year. However, this is true only if extraction and processing costs are zero, or if extraction and processing costs also rise at r per cent a year. Still, if the user costs of a mineral commodity are rising persistently, eventually its price has to rise as well.

However, Hotelling's results depend on a number of strong assumptions. In his model, for example, there is no exploration or additions to reserves, no technological change, and no uncertainty. Modified Hotelling models that relax his assumptions, not surprisingly, no longer make clear predictions regarding user costs and prices. These variables can rise or fall over time. Surveys of this literature are found in Bohi and Toman (1984) and in Krautkraemer (1998).

This means that empirical studies are needed to determine if the influence of technological change, new discoveries, and other developments pushing real mineral

³ Others are real user costs and the real marginal costs of extraction and processing (see Tilton, 2002, ch. 3).

commodity prices down have more than offset the effects of mineral depletion and other factors tending to increase mineral scarcity over time. Many such studies have been carried out,⁴ and generally they find that prices since the 1870s, despite quite large fluctuations in the short run, have followed either constant or (as in the case of copper) falling long-run trends.

3. INFLATION BIAS IN RECENT DECADES

Claims that the official price series in the United States and elsewhere overstate inflation arose over 40 years ago (Stigler, 1961), but had little impact on the collection and reporting of the official U.S. price indexes over the next 35 years. In the mid 1990s, the interest in price indexes re-emerged and the Boskin Commission was set up with a mandate to evaluate and suggest improvements in the CPI. It found that the CPI overestimates inflation by some 1.1 percentage points per year (Boskin *et al.*, 1996). Subsequently, a string of studies of the inflation bias has come forth, applying different estimation technologies (Table 1).

The renewed interest in price indexes has been spurred by a growing realisation of the adverse consequences of incorrectly measuring inflation. As noted by Boskin *et al.* (1998: 3): “Accurately measuring prices and their change, inflation, is central to almost every economic issue. There is virtually no other issue that is so endemic to every field of economics. Some examples include aggregate economic growth and productivity; industry prices and productivity; government taxes and spending programs that are indexed to inflation; budget deficits and debt; monetary policy; real financial returns; real wages; real median incomes and poverty rates; and the competitive performance of economies.”⁵

To the above list one should add the ramifications for deriving real price trends for primary commodities, the focus of this paper, and the related issue of trends in the barter terms of trade for commodity-exporting countries.

⁴ See, for example, Potter and Christy (1962), Barnett and Morse (1963), Manthy (1978), Pindyck (1978, 1999), Smith (1979), Slade (1982, 1988), and Howie (2002).

⁵ Also see Gordon (1990: 8-14). Among the areas listed, the question of whether the slowdown in productivity growth in the United States since the mid 1970s is overestimated has attracted the most attention. See, for example, Jorgenson and Stiroh (1999, 2000), Onliner and Sichel (2000), and Gordon (2000).

3.1. Sources of bias

There are three main reasons why the U.S. and other price indexes overstate inflation:

Substitution bias. In constructing the price series, the Bureau of Labour Statistics (BLS) uses a modified Laspeyres index. In this type of index, the weights of the different items are derived from consumption expenditure surveys from a particular base year. This base-year “basket” is re-estimated infrequently: in the case of consumer goods, only about every tenth year (Abraham, 2003). When there are shifts in relative prices over time, buyers substitute towards items with falling relative prices, signifying the need for a change in the weights, which the fixed-weight Laspeyres index ignores. Boskin *et al.* (1998: 7) finds the use of such an index “extreme, unrealistic, and unnecessary”, and estimates it has caused the CPI to overestimate inflation by 0.4 percentage point per annum in recent decades.

New goods bias. There are at least three flaws in the official CPI and PPI when it comes to handling new products on the market. The first arises because new products are introduced in the indexes with a long lag. A frequently cited example is the cellular phone, which appeared in 1983, but was not included in the CPI until 1998. Other examples include room air conditioners, VCRs, and personal computers, which were all included “years after they were first sold in the market place” (Gordon and Griliches, 1997: 85). As a result, the accumulated price decreases for these products during the intervening years, estimated at around 80 per cent, were not captured in the official price indexes (Boskin *et al.*, 1997).

A second flaw occurs because improved versions of many products are introduced as completely new goods, which means that the quality improvements they reflect are ignored. A third flaw arises because new products are included in the indexes without allowing for the consumer surplus gain (compensating variation) they generate (Hausman, 1997; 2003). The Boskin report makes no attempt to correct for this third source of bias.⁶ The size of the new-product bias arising from the first two sources, it

⁶ Hausman (1997) has suggested a simple, but ingenious method for dealing with the consumer surplus omission. There are other sources of bias in the CPI that are quantitatively small and hence not detailed here, e.g. the switch to discount stores and spurious weights used to aggregate the CPI (Lebow and Rudd, 2003).

estimates to be in the range of 0.20 to 0.30 percentage points per year (Boskin *et al.*, 1997: Table 1).

The pure-quality bias. In the official price indexes, aside from a few recent exceptions noted below, goods are measured in standard units, such as *a* diesel engine or *an* electronic calculator, without any adjustment for improvements in quality. Thanks to such improvements, the consumer gets more for the price paid, both in a quantitative and qualitative sense.⁷ A calculator today has more capacity and many more capabilities, is less energy demanding, has lower repair costs, and lasts longer, than the calculators available some years ago. The Boskin report argues that prices of final goods and services should be divided into two separate components—one for quality improvements and one for pure price changes (inflation). It estimates that the new-product and pure-quality biases together cause the CPI to overestimate inflation by some 0.60 percentage points per year. The pure-quality bias by itself is responsible for about 0.30 - 0.40 percentage points (see above).

The quality issue has indeed become highly visible to all who follow the market for computers and related electronic products. Here quality improvements, as measured for example by capacity, have far outpaced prices, in fact by 15-30 per cent per year according to several estimates for the 1980-1999 period (Jorgenson and Stiroh, 2000). Such developments have helped raise the awareness of quality improvements in other product categories as well.

3.2. The Engle-curve method

The method used by Boskin *et al.* and Gordon for estimating inflation bias has been dubbed “down the trench”, as it is based on very detailed and laborious examinations of individual products. A much simpler approach, the Engel curve method, was introduced a few years ago.

Authors using this method (see Table 1) take the negative relationship between household income and the share of income spent on food (the Engel curve) as the starting point. With raising per capita income, the average share of food in total household

⁷ It is important to note that the quality improvements estimated by Gordon (1990) exclude intangible quality changes, such as better sound in video recorders.

expenditures has declined steadily in the US (from about 27 per cent in the early 1950s to 13 per cent in the mid 1990s). However, there is little reason to expect that the food share of household expenditure *at a given real income level* should change over time, after controlling for changes in the relative price of food, income distribution, the demographic composition of the population, and other relevant variables.

The estimation method can be illustrated with the help of a simple graph depicting Engel curves (F/Y) for two consecutive years (Figure 1). The available studies show a drift of the Engel curve to the left over time, from say $(F/Y)_t$ to $(F/Y)_{t+1}$. That is, for a given level of real household income (as deflated by the official CPI), the food share has declined over time. They conclude that the observed drift is due to overestimation of inflation, and use it to estimate the bias in the CPI. In Figure 1, the bias is given by $\Delta Y/Y_t$.

Figure 1. The Food Engel curve

3.3. Official responses

Results that upset conventional wisdom and established practice tend to receive mixed reviews. This was case with the Boskin report. The first reactions by officials and researchers at the BLS were largely sceptical. Now, most of the recommendations

advanced by the Boskin Commission are officially endorsed. The Commission recommended the use of a superlative Törnquist index, which is similar to the Fischer ideal price index (Boskin *et al.*, 1998). The Törnquist index is based on averaging (geometrically) the weights from the base and the most current year, which reduces the substitution bias inherent in the Laspeyres index.⁸ The BLS introduced the use of a superlative index in 2002, although only for some “strata” (sub-indexes) in the CPI index (Abraham, 2003). The Bureau of Economic Analysis (BEA) has also recently switched to a superlative index (Moulton, 2000).

BLS officials now also agree that previous methods for handling new products lead to upward bias in the CPI (Moulton, 2000; Abraham, 2003), as argued by the Boskin Commission. New practices have been established or are underway. One important change since 2002 is that the weights used to aggregate the CPI indexes are now updated every two years instead of every tenth year (Abraham, 2003).

When it comes to the third bias, pure quality changes, controversy still lingers, both regarding estimation methods and the numbers advanced by the Boskin Commission. The Commission (Boskin *et al.*, 1996) relied mainly on the hedonic method for separating quality-induced price changes from pure inflation. With this method, price changes for a good are regressed against quality changes, as reflected by technical specifications (such as horsepower, fuel consumption, and expected life in the case of automobiles). The share of the price increases accounted for by changes in such technical characteristics is attributed to quality improvements and the remaining share to inflation.

Objections to the Boskin report on this issue have focused on three related points. First, the method it uses for some product categories, it is claimed, is subjective and judgmental. Second, quality can deteriorate as well as improve, and for many goods and services no reliable methods for estimating quality changes exist (Schultze and Mackie, 2002). Third, the official indexes have increasingly taken quality changes into account in

⁸ For further discussions of the choice of indexes, see Boskin *et al.* (1998), Gordon and Griliches (1997), Deaton (1998), Moulton (2000), Schultze and Mackie ((2002) and Abrahams (2003).

recent years. Examples include computer software, automobiles and telecommunication equipment (Stewart and Reed, 1999).⁹

Boskin (2000) maintains that the Commission's original assessments of quality improvement hold, and that the recent inclusion into the official statistics of quality changes for a few products is insufficient. Jorgenson and Stiroh (2000) also identify further adjustment for quality changes as one of the two top priorities for the agencies producing and using price series. Outsiders are split between those who are sceptical of the estimates from the Boskin Commission (Deaton, 1998; Schultze and Mackie, 2002) and those who find them too conservative (Diewert, 1998; Nordhaus, 1997; Hausman, 2003).¹⁰

In the past few years, researcher at the BLS have themselves made some simulation of bias in the CPI during the past two decades. In one of these studies, Stewart and Reed (1999) carries out a backward re-estimation of "what would have been the measured rate of inflation from 1978 forward had the methods currently [1999] used in calculating CPI-U[rb] been in use since 1978". These new practices (14 in all) concern mainly quality change adjustment for electronic products (including computers and television sets) and a geometric formula that "assumes a modest degree of consumer substitution within most CPI item categories." They find that annual inflation would have been 0.45 points lower had these practises been applied throughout the entire 1978-1999 period. Abraham (2003), who served as a commissioner at the BLS up to 2001, made another simulation: how would the 2002 new practice of introducing new goods more expediently have affected the CPI in earlier years? She finds that the CPI would have been 0.17 points lower per annum since 1987.

Adding up the overestimation of CPI due to neglect of quality improvements and consumer substitution, and late introduction of new goods, as estimated by Stewart and Reed respective Abraham, suggests that inflation in recent decades has been overestimated by at least 0.62 percentage points (0.45 + 0.17). This number is not that

⁹ The methods used by the BLS to adjust for quality changes are described in Moulton and Moses (1997). Moulton (2000) acknowledges that quality change is insufficiently incorporated into the official indexes, that improvement is warranted, and that this is underway.

¹⁰ The main exchanges in this debate can be consulted in the May issues (Papers and Proceedings) of the *American Economic Review* for 1997, 1999 and 2000; the *Brookings Papers on Economic Activity*, 1997:1;

different from the low-point estimate derived by Boskin et al. (0.8 percentage points). Moreover, it should be noted that the BLS revisions are understated since they take into account only quality changes for a few selected products. The Boskin Commission claims that quality improvements in a long range of other products should also be accounted for.¹¹

Revisions in the official CPI indexes since 1983 probably explain why many studies listed in Table 1 (Gordon, Hamilton and Costa) report a decline in bias over time. This may also be the reason why Lebow and Rudd (2003) found the overall, and the quality, bias for the recent period 1987-2001 to be somewhat smaller than reported by Boskin *et al.* (1996) for 1974-95 (Table 1).

In summary, there is now wide agreement that the CPI has overestimated inflation for the four decades prior to the mid 1990s, when more important revisions were initiated. Disagreement and uncertainty remain over the size of the bias (as is evident from the large confidence intervals reported by most authors in Table 1). On the methodological front, the debate is now largely focused on how to best account for quality change (Schultze, 2003). It is important, finally, to recall that the historical CPI index has not been revised (Stewart and Reed, 1999). So all the biases reported for recent decades exist in the price indexes for earlier years. This is particularly important for our purposes, since almost all of the period we are examining is before the 1990s.

4. INFLATION BIAS BEFORE 1950

There are two main hurdles when it comes to estimating the inflation bias over the 1870-1950 period. First, while there is one systematic attempt to estimate inflation bias for years prior to 1950 (Costa, 2001), it does not take into account quality change. Therefore, we have to rely on evidence of quality change from numerous case studies. Second, official CPI index was first compiled in the US in 1917 and stretched back to 1913.

and the *Journal of Economic Perspectives*, Winter 1998 and Spring 2003. To our knowledge, Boskin and his co-authors have not yet responded to the Schultze and Mackie report (2002).

¹¹ According to Bosworth (1997), the quality adjustment made by the BLS lowered the official CPI by between 0.28 and 1.76 percentage points (with a mean of about 1.0) for the year 1995. However, these quality adjustments are for selected goods only and the average is derived on the assumption that all other products had no quality change. This introduces an overall downward bias of unknown size. This estimate nevertheless suggests that unadjusted quality changes in the years before the early 1990s were notable, providing further support for the quality bias estimates of the Boskin Commission.

Earlier estimates “are based on the efforts of economists and economic historians who have created CPI estimates and linked them to the modern series” (McCusker, 2001). This raises questions about the reliability of inflation figures before 1913 (Sahr, 2002).

4.1. Substitution bias and new goods bias

As the BLS has used the Laspeyres since 1917, the substitution bias that it produces clearly dates back to the earliest CPI figures. Similarly, the late introduction of new goods is not a new problem. It seems to have been as pronounced, if not more so, before the 1950s as after. Stigler *et al.* (1961: 38) cites as examples that “new automobiles were not introduced until 1940, used automobiles.... until 1952”, some 40-50 years after they were on the market, and that “refrigerators were introduced only in 1934”, probably with an equally long lag. They conclude that: “Our observations suggest that in many cases the need to introduce a new variety of a commodity is felt [by the BLS] only when it becomes impossible to price the old variety because it has disappeared from the outlets”. (Rather the same observation is made by Abraham (2003: 52), reflecting on the 1990s.)

The only comprehensive study that has estimated inflation bias before 1950 is Costa (2001). The Engel-curve method she uses takes all technical index-related flaws into account, but not the pure quality bias. As Table 1 shows, her partial estimate indicates practically no inflation bias in the 1888-1919 period, and a bias in the range of 0.4 to 1.1 percentage points for the 1919-1935 period. Her mid-point estimate of the substitution and new goods biases for this period (at 0.7 points), is of the same order as reported by the Boskin Commission for the 1975-95 period (about 0.6 points; cf. pp. 6-7 above).

4.2. Pure-quality bias

The type of systematic and broad-based estimations of quality change in products and services that the studies cited in Table 1 have undertaken for recent decades does not exist for the period before 1950. The study by Costa, as noted earlier, does not consider quality improvements. This, however, does not mean that information is totally lacking. For instance, Griliches (1961) estimates that quality improvements accounted for more than half of the price increase that the U.S. Wholesale Price Index (WPI) reported for

automobiles over the 1937-1960 period.¹² So less than half the increase was pure inflation. Gordon (1990: 11) cites a study by Mills from 1936, dealing with agricultural machinery and motor cars, that found substantive quality improvements over the 1910-32 period that the WPI had failed to consider. Another study by Rees revealed rapid quality improvements in rails and railway equipment between 1890 and 1914 (Gordon, 1990). There is also the celebrated history of lighting by Nordhaus (1997), the work by Raff and Trajtenberg (1997), again on automobiles, and some others referred to in Bresnahan and Gordon (1997) and Stigler *et al.* (1961). The quantitative estimates of the quality change in the particular goods examined in all these case studies are similar to, or greater than, those reported by the studies listed in Table 1 for later periods.

It is also the case that during the 1920s and 1930s, a long range of what were then technically sophisticated consumer goods, including telephones, radios, and electrical appliances of various types, became widely available for the first time. Clearly there were many quality improvements in such products in the years up to 1950, although they have not been estimated.

There is also an *a priori* reason to expect that numerous major technological innovations in the past—including electricity, railroad equipment, the internal combustion engine, a wide range of petrochemicals, and communication equipment—led to major quality improvements in many goods with such inputs during the first half of the 20th century. Gordon (1990: 11-12) conjectures that one would “suspect that existing capital goods deflators significantly overstate quality-corrected inflation in capital goods prices . . . all the way back to the dawn of the industrial age”. Nordhaus (1997) arrives at a similar conclusion and provides an exhaustive list of “great inventions” dating back to the 19th century, whose influence on prices is not reflected in the official indexes.

While there is little reason to believe that quality improvements in technically advanced products were any less pronounced before 1950 than after, the overall bias induced by quality changes may nevertheless have been smaller. Far less quality change tend to occur in what Nordhaus (1997) labels as “run-of-the-mill sectors”. According to his estimates, the importance of these sectors has declined from 75 to 28 per cent of the

¹² The U.S. Wholesale Price Index became the U.S. Producer Price Index in 1978. It is also notable that the WPI before 1947 contained almost no producer goods (Gordon, 1990: 15).

U.S. economy between 1800 and 1991. Along with this shift, from standard bulk goods to more technologically sophisticated products, the overall quality bias may have grown. Although no quantitative data are provided, Gordon (1999: 126) in extensions of his previous work on durables, found “much smaller rates of [quality] bias before 1947”. This conjecture is shared by Boskin *et al.* (1996: 42) for the CPI.

5. COPPER PRICE TRENDS

This section estimates the trend in real copper prices over the 1870-2000 period using as deflator an adjusted CPI that reduces the reported inflation. In the benchmark case, we take the inflation bias to be 1.0 percentage points a year throughout the period. This number is smaller than the estimates for recent decades reported by most of the studies listed in Table 1, and slightly above the estimate by Costa for the 1919-35 period (which disregards quality change). It is also the number that Nordhaus (1997) comes up with in his “thought-experiment” (*gedankenexperiment*). In addition to this benchmark or base case, we subsequently provide alternative estimates of the real price trend for copper, assuming that the CPI bias is 0.5 and 1.5 percentage points. Our objective is not to advocate one particular figure, but to consider a range within which the true bias is likely to fall. We can then explore how the long-run trend in the real price of copper changes as the magnitude of the bias varies over the range of likely possibilities.

Figure 2 shows the average annual U.S. producer price of copper for the years 1870 to 2000 deflated by the CPI minus 0.0, 0.5, 1.0, and 1.5 percentage points a year. When no adjustment is made to the CPI, a downward trend is quite apparent. As the annual adjustment to the CPI increases from 0.0 to 1.5 percentage points, a reversal in trend appears to occur, calling into question the presumed decline in real copper prices. The following sections examine the statistical significance, stationarity, and robustness of these trends. It begins with the base case and then examines how the trends change when the assumptions are altered

5.1. The base case

The base case focuses on the real copper prices over the period 1870 to 2000. It considers the average annual U.S. producer price, in the tradition of Potter and Christy (1962),

Figure 2. The U.S. Producer Price of Copper Deflated by the CPI from 1870 to 2000 with 1950 = 100

Barnett and Morse (1963), Manthy (1978), and Howie (2002), rather than the comparable LME price of copper. It uses the U.S. CPI to deflate nominal prices. This is in part because this index goes back to 1870 (although the index for 1870-1913, as noted earlier, is not an official index). Moreover, the inflation bias in the CPI has been studied more thoroughly than in other price indices. Of greatest importance, however, is the fact that the CPI reflects the real price of copper in terms of a representative basket of consumer goods and services. As a result, it more accurately assesses the impact of commodity price trends on the welfare of society—that is, consumers or people in general—than other deflators.

The base case further assumes that the long-run trend in real copper prices is linear and it reduces, as mentioned, the official inflation rate by one percentage point a

year to correct for the overestimation of the CPI. Specifically, the base case estimates the following equation:

$$P_t = \beta_0 + \beta_1 t_t + \varepsilon_t \quad (1)$$

P_t is the average U.S. producer price of copper for year t , deflated by the CPI adjusted by subtracting one percentage point each year from the annual increase in the CPI. The variable t_t is the time trend. It takes the value of 1 for 1870 (or for 1871 or 1872 when lagged values of the dependent variable are included on the right hand side of the equation to eliminate autocorrelation) and increases by 1 for each succeeding year. The β 's are the parameters to be estimated and ε_t is disturbance term. The results for this model are compared with the results when no adjustment is made to the CPI.

Since the independent variables (time and the constant) are predetermined, we use generalized least squares (GLS) to estimate the parameters. We also use the augmented Dickey-Fuller (ADF) procedure—or where necessary the Ouliaris, Park, and Phillips (OPP) test—to test for stationarity, as only data series that are trend stationary can be modelled in their level forms (that is, undifferenced) without generating possibly spurious results.¹³

Table 2 shows that the long-run trend in real copper prices derived with the unadjusted deflator is downward. This decline is statistically significant at the 1 per cent level. In contrast, the results for the base case (adjustment by 1.0 percentage point) suggest that the real price of copper has trended upward over the past 130 years. Though significant only at the 12 per cent level, the estimated parameter for the time variable is positive.

¹³ A stationary variable has a constant mean and variance over time. A trend stationary variable has a constant mean and variance once its trend is removed. A differenced stationary variable has a constant mean and variance in its first differences. A variable that is not trend stationary may not follow the same trend over time. This means that the true β 's in equations 1, 2, and 3 may not be constants, but instead may vary over the period examined. In this case, the estimated β 's reflect some hybrid average of the true β 's, and as a result are unlikely at any point in time to represent the true long-run trend in prices.

5.2. Robustness tests

(1) *Alternative Adjustment Factors.* The results when we re-estimate equation 1, using annual deductions of 0.5 and 1.5 percentage points from the annual increase in the CPI, are also shown in Table 2. The lowest estimate of the bias (0.5 points) can be thought of as representing a case where no quality improvements in goods and services occurred over the entire 1870-2000 period. In this case, the only biases in the CPI stem from the use of the fixed-weight Laspeyres index and the delayed introduction of new goods. According to Stigler (1961), these flaws plagued the official CPI index from the time it was first introduced in 1917 and all the way to 1999, when it was first modified (Stewart and Reed, 1999).

The second alternative, a correction of 1.5 percentage points per year, is similar to the upper tail of the range of bias estimates put forth by Boskin *et al.* (1996). It also squares with the mid-point estimates derived by Gordon, Hamilton, and Costa for years in the post-1950 period. It is worth noting that the latter two studies do not take quality change into consideration. The one study based on the Engel curve method that does (Bils and Klenow, 2002) reports a higher number for most consumer durables (2.2 points).

An adjustment of 0.5 percentage points suggests that the long-run price trend is downward as when no adjustment is made to the CPI. However, it is no longer significant. With the 1.5 percentage point deduction, the long-run trend is upward, as in the base case, and statistically significant at the 1 per cent level.

(2) *Alternative Copper Prices.* Using the LME price instead of the U.S. producer price does not affect the results, as Table 3 shows. When no adjustment is made to the CPI, the long-run trend in the real LME price of copper is downward and significant at the 1 per cent level. When the CPI is reduced by 1.0 percentage points a year, the long-run trend is upward, and the coefficient on the trend variable is significant at the 6 per cent level. However, the ADF test in this case suggests that the LME price of copper may not be trend stationary, raising the possibility of spurious results.

(3) *Alternative Deflator.* Several earlier studies use the producer price index (PPI) to deflate nominal commodity prices. To determine if our results are sensitive to the choice of deflator, we re-estimated equation 1 using the PPI deflator.¹⁴ Copper prices deflated by both the adjusted and unadjusted PPI are trend stationary. When the PPI is reduced by one percentage point a year, the long-run trend in real copper prices is upward, as in the base case, though it is now significant at the 1 per cent level. When no adjustment is made to the PPI, the estimated trend is significantly negative. So using the PPI instead of the CPI changes our findings very little.¹⁵

(4) *The 1950-2000 period.* To determine how sensitive the findings are to the length of the period considered, we also re-estimated equation 1 for the years 1950 to 2000. The results are shown in Table 4, and again indicate a significant downward trend when no adjustment is made to the CPI. When the CPI is reduced by one percentage point a year, the coefficient on the time variable is negative, not positive as with our previous results. However, the estimated time-coefficient is insignificant and the ADF test fails to reject the hypothesis that the series is difference stationary, signifying that the results may be spurious.

(5) *Excluding Non-competitive Market Periods.* According to Herfindahl (1959) and Howie (2002), interventions by governments and collusive actions by producers distorted the market price for copper during the following periods: 1870-83, 1888-90, 1900-01, 1914-18, 1929-35, 1939-57, and 1962-64. There are (at least) two, quite different, ways to take account of such market manipulations. The choice depends on how one believes temporary market interference affects long term prices.

One possibility is that the long-term price trend is unaffected, on the grounds that temporary price spikes occurring during periods of market interference are entirely offset during subsequent periods. So once they are over, a competitive market for some time generates prices below the level that would otherwise have prevailed. Successful efforts to raise copper prices artificially reduce demand and stimulate supply. Just the opposite happens when price controls or other market interventions keep the price below its

¹⁴ It is worth noting that Gordon (1990) found the bias in the price index for producer durables to be much higher than for consumer durables (3.0 versus 1.5 percentage points, respectively) over the 1947-1983 period.

competitive level. Demand grows more rapidly and supply more slowly. Once a competitive market is re-established, price is higher for a time than it would have been. Under these assumptions, nothing has to be done in the estimation of long-term price trend to account for market interventions. We then “only” have the problem that the artificial price swings caused by market interference, induce noise in the estimations, which lower the t-values and the overall fit, but have no effect on the coefficients to be estimated (the β ’s).

The alternative (extreme) assumption is that once the supply restrictions are relaxed, prices move directly back to their long-run competitive trend. That is, artificial price volatility has no effect on the long-term trends in either supply and demand. In this case we can estimate the true long-run price trend during periods under competitive conditions by omitting the years when collusion is in effect. Table 4 shows the results from such estimations when the CPI is adjusted by 0.0 and 1.0 percentage points a year. As before, the coefficient on the time variable when no adjustment is made to the CPI, is negative and significant. When the CPI is adjusted by 1.0 percentage points, the coefficient is positive, but insignificant. Furthermore, the ADF test fails to reject the hypothesis that the series is difference stationary, raising the possibility of spurious results.

(6) *Alternative Trends.* The base case assumes that the long-run trend in real copper prices is linear. There are, of course, other possibilities. Slade (1982), for example, argues the trend is quadratic, falling for a time and then rising. Yet another alternative is the inverse function, which has the advantage over the linear trend that it need not at some point become negative when prices are declining. It is more difficult to rationalise, however, if real prices are rising. To assess how dependent our results are on the assumption of a linear price trend, we estimated these two non-linear functions—the quadratic and the inverse—using equations 2 and 3.

$$P_t = \beta_2 + \beta_3 t_t + \beta_4 t_t^2 + \varepsilon_t \quad (2)$$

¹⁵ We also tried the GDP deflator, but it does not produce a copper price series that is stationary, perhaps because it goes back only to 1929 and so generates a shorter series.

$$P_t = \beta_5 + \beta_6 t_t^{-1} + \varepsilon_t \quad (3)$$

Since the ADF test is valid only when the underlying trend is linear, we used the Ouliaris, Park, and Phillips (OPP) procedure to test for stationarity. When copper prices are deflated by the adjusted CPI (using a 1.0 percentage point annual reduction) and the unadjusted CPI, this test indicates that the resulting real copper price series are trend stationary. This finding holds both when real copper prices are assumed to be following a quadratic trend and an inverse trend.

The Schwartz Criterion, which is conventionally used to compare the performance of different equation specifications, indicates that the inverse relationship provides a slightly better fit than the linear trend, and that both the inverse and linear trends do better than the quadratic. The inverse function, however, implies that the price of copper is gradually approaching some stationary or permanent level, which seems questionable when prices are rising

Whether one uses the linear or the inverse relationship, however, does not greatly alter our earlier findings. With the inverse function, the trend in real copper prices is significantly downward if no adjustment to the CPI deflator is made. When the CPI is reduced by 1 percentage point a year, the trend is upward, and the coefficient is statistically significant at the 10 per cent level.

Of course, if the quadratic were the true long-run trend, this would have important implications for availability and prices in the future. The estimated quadratic trends reported in Table 4 indicate a decline in the long-run trend in copper prices until 1986 when no adjustment is made to the CPI and a decline only until 1890 when the CPI inflation rate is reduced by a percentage point a year. Thereafter, the indicated trends rise. There is, however, little to suggest the true long-run trend is quadratic. The coefficients for the square of the time variable, as shown in Table 4, are insignificant, and as noted earlier, the Schwartz criterion favours both the linear and inverse over the quadratic relationship.

6. CONCLUSIONS

Growing evidence indicates that the US CPI and other standard deflators overestimate inflation. The magnitude of the bias is uncertain; the estimates range from 0.9 per cent to over 2.0 per cent a year, and vary somewhat with the time period covered. Though most of the evidence applies to the period after World War II, there are good reasons to believe that real price series for earlier years, even back to the 1870s, are biased as well.

When the nominal price of copper is converted to a real price using the CPI with no correction for the overestimation of inflation, the linear trend over the past 130 years is downward and statistically significant. However, when the CPI is reduced by 0.5, 1.0, and 1.5 percentage points a year to correct for its upward bias, this finding is called into question. With the two higher adjustments, the estimated trend is upward and with the 1.5 percentage point adjustment, it is statistically significant at the 1 per cent level.

Moreover, this finding is quite robust. It holds whether nominal copper prices are those reported by the North American producers or the LME. It holds if the PPI is used to deflate nominal prices instead of the CPI. It holds if one assumes the true trend in real prices follows an inverse relationship rather than a linear one. It does not hold, however, when periods of market manipulation are excluded from the sample. The estimated time coefficient then becomes insignificant and the stationarity test indicates that the results could be spurious. It is worth noting, though, that with 1.0 point adjustment, the trend coefficient is not negative.

These results derived here do presume that the price of copper has followed one constant secular trend over the past 130 years. While tests for trend stationarity, as we have seen, provide some support for this assumption, there are reasons to suspect that the long-run trend in copper prices has changed from time to time as important new end uses for copper have arisen, as close substitutes have been developed, as major innovations have reduced copper production costs, and as the world economy itself has experienced structural change. Empirical evidence indicating that copper prices are not trend stationary over the shorter 1950-2000 period, or over the longer 1870-2000 period when the years of market manipulations are excluded, adds some support for this possibility.

If in fact the long-run trend in copper prices has changed from time to time, this provides yet another reason to question the widespread belief that the price of copper has

persistently for a century or longer been following a downward trend. In any case, whether the long-run trend has or has not been following a single long-run trend does not alter the fact that the actual path of real copper prices properly deflated for inflation has remained constant or risen (depending on the assumed size of the inflation bias), rather than fallen, over the past 130 years.

To what extent this finding for copper can be generalized awaits similar studies on other mineral commodities. Comparing copper price trends as conventionally estimated, however, with those for other mineral commodities provides some tentative indications. The real price of copper without any adjustment to the deflator, as we have seen, falls over the long run. In this respect it is similar to a group of other mineral commodities, including aluminium, nickel, and silver. Here the impact of adjusting the deflator for other mineral commodities is difficult to determine in advance. In the case of copper, we have seen that the long-run trends in real prices are constant or rising once the deflator is corrected for inflation bias. However, in the case of aluminium, nickel, and silver, the decline in the unadjusted price series is greater than for copper, and may be sufficiently large to persist even after adjusting the deflator.

Another group of mineral commodities displays an unadjusted long-run price trend that is more or less constant, neither rising nor falling significantly. This group includes petroleum, lead, zinc, and tin. Adjusting the deflator for these commodities, as we have done in the case of copper, would probably convert the stationary trend in real prices into a rising trend (although short-term fluctuations are quite pronounced).

It is important, however, to keep these results in perspective. Even if mineral scarcity is rising somewhat, what is truly surprising is that the price increases have not been greater. Given the sharp growth in population and the increase in per capita incomes particularly in the developed countries, the consumption of all mineral commodities over the past 130 years has exploded. Despite the huge increases in production that have made this possible,¹⁶ the real prices of mineral commodities have not increased greatly. Moreover, the price increases that may have occurred have clearly not prevented the world economy from growing strongly over the past century.

¹⁶ World copper mine output, for example, rose from under 0.5 million tons in 1900 to 13.2 million tons in 2000 (Porter and Edelstein, 2003).

The possibility that the real prices for many mineral commodities have been rising rather than falling over the past century has important implications for development economics as well as resource economics. In particular, the barter terms of trade for those countries that export primary commodities and import manufacturing products may have declined less than the available estimates suggest, or may in many cases not have declined at all. This is because the price trends for primary commodities, as measured by the standard price indexes, are largely free of the biases that beset the price trends of imported manufactured goods. For primary commodities, there is much less of an issue with quality change and the delayed introduction of new products. To what extent the upward bias in the prices of imported manufactured goods has produced a downward bias in the barter terms of trade for countries exporting primary commodities is another area that this study suggests is in need of attention.

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**Table 1: Estimated Bias in the US Consumer Price Index, Various Studies and Methods
(percentage point bias per annum)**

Study (year)	Coverage and special features	Period(s) covered	CPI bias (%)	Range of CPI bias (%) Low High	
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a) Studies based on Hedonic and related estimation methods

Gordon (1990)	About half of all consumer durables. Low estimate assuming no quality change in non-covered items; high estimate same for all items	1947-1960	2.2	2.2	4.4
		1960-1972	1.2	1.2	2.5
		1972-1983	1.1	1.1	2.1
		1947-1983	1.5	1.5	3.1
Boskin <i>et al.</i> (1996)	Quality bias for selected items from 27 main categories in CPI. Extrapolation to other products	1975-1995	1.1	0.8	1.6
Nordhaus (1997)	“Thought-experiment” (Gedankenexperiment)	1800-1992	1.0	0.5	1.4
Lebow (2001)	Re-estimation of Boskin <i>et al.</i> results for a later period	1987-2001	0.9	0.3	1.4

b) Studies based on the Engel-curve method

Hamilton (2001)	1,800 households and no account of pure quality change. Food as the instrument good	1974-1981	2.5	2.5	2.9
		1981-1991	1.0	1.0	1.7
		1974-1991	1.6	1.6	2.2
Costa (2001)	Income and food expenditure data for 14,000 to 26,000 households. No account of pure quality change. Food and recreation as instrument goods. The 1888-90 survey is not representa- tive (heavy industry workers only). CPI (=food) for 1888-1919 unreliable and unofficial estimate	1888-1919	-0.1	-0.1	-0.3
		1919-1935	0.7	0.4	1.1
		1935-1960	...	...	...
		1960-1972	0.4	-0.2	1.1
		1972-1982	2.7	1.9	3.9
		1982-1994	0.6	0.1	1.6
		1972-1994	1.6	1.4	2.6
Bils and Klenov (2001)	Quality-adjusted Engel curves for 66 consumer durable goods (80% of all and 12% of CPI)	1960-1994	1.1	1.1	1.3
		1980-1996	2.2	0.8	2.4

c) Simulations of selected biases conducted by BLS officials^{a)}

Bosworth (1997)	Quality bias for selected goods in one year only	1995	1.0 ^{b)}	0.3	1.8
Stewart and Reed (1999)	Estimates of what would have been the CPI if 1999 estimation practices had been applied 1978-1998	1978-1998	0.45	-	-
Abraham (2003)	Updating weights more frequently than actually done and geometric formula	1988-1999	0.37	-	-

Sources: See reference list

a) These studies estimate only part of the total bias in the CPI.

b) Only the range is reported in this study; the average is derived from this range.

Table 2. Results for the U.S. Producer Price of Copper Deflated by CPI and Adjusted CPI^a

	CPI Adjustment			
	Conventional (0.0 %)	Low (0.5 %)	Base Case (1.0 %)	High (1.5 %)
Constant				
Coefficient	195	136	89	50
t-statistic	8.50	8.27	5.32	3.23
Probability	0.00	0.00	0.00	0.00
Time				
Coefficient	-0.93	-0.25	0.34	0.87
t-statistic	-3.76	-1.18	1.58	3.62
Probability	0.00	0.24	0.12	0.00
ADF test statistic ^b	-4.03	-4.40	-4.11	-3.85
Adjusted R-squared	0.79	0.67	0.72	0.84
Durbin-Watson statistic	1.89	1.90	1.91	1.93
F-statistic	159	88	108	223
Schwarz criterion	9.43	9.11	8.93	8.89

Notes: ^aAll equations are estimated using GLS with a two-lag autoregressive residual. Some t-statistics are derived from White heteroscedasticity-consistent covariance matrix.

^bThe critical ADF test statistic values are -4.03, -3.45, and -3.15 for significance at the 1 per cent, 5 per cent, and 10 per cent levels respectively.

Table 3. Results (a) for LME Copper Prices Deflated by the CPI and (b) for U.S. Producer Prices Deflated by the PPI^a

	LME Price Deflated by the CPI		U.S. Producer Price Deflated by the PPI	
	CPI Adjustment (0.0%) (1.0%)		PPI Adjustment (0.0%) (1.0%)	
Constant				
Coefficient	174	79	189	78
t-statistic	8.54	5.44	8.69	4.83
Probability	0.00	0.00	0.00	0.00
Time				
Coefficient	-0.71	0.44	-0.73	0.66
t-statistic	-2.73	1.91	-2.98	2.83
Probability	0.01	0.06	0.00	0.01
ADF test statistic ^b	-4.32	-2.77	-3.98	-3.59
Adjusted R-squared	0.73	0.71	0.75	0.79
Durbin-Watson statistic	1.93	1.90	1.93	1.96
F-statistic	116	106	127	151
Schwarz criterion	9.52	9.37	9.29	8.91

Notes: ^aAll equations are estimated using GLS with a two-lag autoregressive residual. Some t-statistics are derived from White heteroscedasticity-consistent covariance matrix.

^bThe critical ADF test statistic values are -4.03, -3.45, and -3.15 for significance at the 1 per cent, 5 per cent, and 10 per cent levels respectively.

Table 4. Results for the U.S. Producer Price of Copper Deflated by CPI and Adjusted CPI (a) for 1950-2000 and (b) for the 1870-2000 Excluding Non-Competitive Market Periods^a

	1950-2000		1870-2000 Excluding Non-Competitive Market Periods	
	CPI Adjustment (0.0%)	(1.0%)	CPI Adjustment (0.0%)	(1.0%)
Constant				
Coefficient	158	171	209	108
t-statistic	7.99	6.02	4.57	2.55
Probability	0.00	000	0.00	0.01
Time				
Coefficient	-1.71	-1.08	-1.00	0.17
t-statistic	-2.81	-1.25	-2.09	0.38
Probability	0.01	0.22	0.04	0.71
ADF test statistic ^b	-2.90	-2.68	-2.01	-2.07
Adjusted R-squared	0.71	0.57	0.76	0.76
Durbin-Watson statistic	1.76	1.76	1.67	1.84
F-statistic	60	33	111	109
Schwarz criterion	8.62	9.11	9.02	8.92

Notes: ^aAll equations are estimated using GLS with a two-lag autoregressive residual. Some t-statistics are derived from White heteroscedasticity-consistent covariance matrix.

^bThe critical ADF test statistic values are -4.15, -3.50, and -3.18 for significance at the 1 per cent, 5 per cent, and 10 per cent levels respectively for the 1950-2000 period. The comparable values for the 1870-2000 period excluding non-competitive market periods are -4.09, -3.47, and -3.16.

Table 5. Results for U.S. Producer Copper Prices Deflated by the CPI Assuming Quadratic and Inverse Trends^a

	U.S. Producer Price Deflated by the CPI, Quadratic Trend		U.S. Producer Price Deflated by the CPI, Inverse Trend	
	CPI Adjustment (0.0%) (1.0%)		CPI Adjustment (0.0%) (1.0%)	
Constant				
Coefficient	230	101	93	131
t-statistic	6.34	3.73	8.44	8.87
Probability	0.00	0.00	0.00	0.00
Time Variable				
Coefficient	-2.37	-0.12		
t-statistic	-2.16	-0.13		
Probability	0.03	0.90		
Time Variable Squared				
Coefficient	0.01	0.00		
t-statistic	1.45	0.52		
Probability	0.15	0.60		
Inverse of Time Variable				
Coefficient			1487	-992
t-statistic			3.15	-1.71
Probability			0.00	0.09
OPP test statistic ^b	23.21	16.81	25.19	14.64
Adjusted R-squared	0.79	0.71	0.80	0.72
Durbin-Watson statistic	1.88	1.91	1.96	1.78
F-statistic	122	81	175	109
Schwarz criterion	9.44	8.97	9.35	8.93

Notes: ^aAll equations are estimated using GLS with a two-lag autoregressive residual. Some t-statistics are derived from White heteroscedasticity-consistent covariance matrix.

^bThe critical OPP test statistic values are 19.95, 15.61, and 13.49 for significance at the 1 per cent, 5 per cent, and 10 per cent levels respectively.

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