

Industrialization and the demand for mineral commodities[☆]

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ABSTRACT

This paper uses a new data set that begins in 1840 to investigate how industrialization affects the derived demand for mineral commodities. I establish that there is substantial heterogeneity in the long-run effect of manufacturing output on demand across five commodities. A one percent increase in per capita manufacturing output leads to an approximately 1.5 percent increase in aluminum demand and a roughly 1 percent rise in copper demand. Estimated elasticities for lead, tin, and zinc are below unity. My results suggest that the experience of Japan and South Korea's industrialization, for example, may be used to infer the impact of China's industrialization on future demand for metals. The results imply substantial differences across commodities with regard to future demand. Adjustment to equilibrium takes 7–13 years, which helps explain the long duration of commodity price fluctuations.

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1. Introduction

The booms and busts in the prices of commodities, such as crude oil and metals, strongly affect the macroeconomic and fiscal conditions of commodities exporting and importing countries (see e.g. [Bernanke, 2006](#); [IMF, 2012](#)). These effects are especially important in developing countries, which rely on exports of a rather narrow set of commodities (see [Van der Ploeg, 2011](#), for a survey).

[Kilian \(2009\)](#) and [Stuermer \(forthcoming\)](#) show that these boom and bust periods are primarily driven by global demand shocks. For example, China's rapid industrialization and its recent slowdown strongly affect world commodity prices. Thus, understanding how industrialization affects the derived demand for mineral commodities is important for macroeconomic and fiscal policy making in commodity exporting developing countries. Given that background, this paper poses the following questions: How does a change in manufacturing output affect the quantity demanded of mineral commodities? What is the price elasticity of demand? Can we utilize experience from past periods of industrialization, e.g. in Germany or Japan, to infer the impact of China's industrialization on the demand for metals?

Empirical evidence on the nexus of industrialization and the derived demand for mineral commodities remains limited. Studies on the elasticities of demand with regard to manufacturing output and prices cover only relatively short periods

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(see [Hamilton, 2009](#); [Pei and Tilton, 1999](#); [Kilian and Murphy, 2014](#), for surveys of the current literature). They typically neither capture the effects of common technological and other factors across countries nor take into account long-term adjustment processes of particular importance in the mineral commodities sector (see [Radetzki, 2008](#); [Stuermer, forthcoming](#)).

The literature also rarely addresses the extent of parameter commonality across countries. One would expect that long-run equilibrium relationships in base-metal processing manufacturing are similar across countries, as these markets have been highly integrated over a long time. At the same time, it is less compelling to assume that short-run relationships should be the same. For example, adjusting production capacity of metals-based manufacturing products might differ in the short run due to different labor and capital market frictions across countries.

This paper explores the link between industrialization and the derived demand for mineral commodities based on a new unbalanced panel data set for a period partially going back to 1840. I assemble a new annual data set for 15 countries, covering by country real manufacturing output and by country real prices and consumption of five non-renewable resources—aluminum, copper, lead, tin, and zinc. These five base metals have characteristics, such as a substantial track record of industrial use and integrated world markets, which make a long-run analysis possible.

My estimation strategy relies on an extension of the partial adjustment model, in which I introduce homogeneity of parameters in a stepwise manner following [Pesaran et al. \(1999\)](#). This allows me to stay a priori agnostic about the commonality of coefficients for the short-term and long-term relationships. I also control for common trends and time fixed effects in a stepwise manner. This allows me to take advantage of the data's panel structure and to control for a variety of omitted common factors such as technological change in resource efficiency or world wars, which might affect demand in all countries at the same time.

I find that the long-run elasticity of metal demand with respect to manufacturing output is very similar across 15 countries, while there is substantial heterogeneity in the short-term coefficients. This suggests that one can utilize past industrialization experiences to infer the impact of China's industrialization on metals demand.

Across five examined commodities, I find substantial heterogeneity in the long-run effect of a change in per capita manufacturing output on the per capita quantity demanded of mineral commodities. A 1 percent increase in manufacturing output leads to an approximately 1.5 percent increase in aluminum demand and a roughly 1 percent rise in copper demand in the long run. Estimated elasticities for lead, tin, and zinc are far below unity. Holding all other factors constant, the intensity of use¹ of aluminum in the manufacturing sector increases over the course of industrialization, while the intensity of copper use is constant, and the intensities of lead, tin, and zinc use decrease. Common linear time trends only have a significant negative effect on the demand for lead and zinc.

Heterogeneity in the effect of manufacturing output on the demand for mineral commodities implies large differences in the amplitude of demand shocks on the prices across the examined commodities. For example, an unexpected slowdown in the growth rate of Chinese manufacturing output will have a stronger negative effect on the demand for aluminum or copper than on the demand for zinc, tin, or lead. This observed heterogeneity may drive differences in the relative contribution of demand shocks on real prices, as found by [Stuermer \(forthcoming\)](#), and in overall price volatility across commodities. I find slow speeds of adjustments of 8–15 percent per year, which imply that it takes about 7–13 years for these markets to reach equilibrium after a shock. The lead market is slowest to adjust, while the tin and copper markets return to equilibrium fastest. This helps explain the longitude of price fluctuations in these markets (see also [Slade, 1991](#), who first introduced this argument to the literature).

The estimated long-run price elasticities of demand are rather inelastic for the examined mineral commodities. Again, there are pronounced differences across the examined mineral commodities. While price elasticity is about -0.7 in the case of aluminum, it is about -0.4 for copper demand, and below or equal to about -0.2 for lead, tin and zinc demand. This shows that these mineral commodities are rather essential to manufacturing output, as the processing industry changes its use slowly in response to price.

Based on my results, countries dependent on mineral commodity exports may better judge the long-term perspective of the respective markets and adjust their macroeconomic and fiscal policies accordingly. For example, the estimated manufacturing output elasticities of demand suggest that industrialization in China will cause aluminum demand to increase relative to manufacturing output, while copper will grow in proportion to manufacturing output. The demand for lead, tin, and zinc decreases relative to manufacturing output in the long term. My results also help firms in the extractive sector define their long-term investment strategies and, hence, facilitate smoother markets.

Current theoretical models of the long-run demand for non-renewable resources do not account for the heterogeneity across resources in the elasticity of demand with regard to intermediate goods or aggregate output. Future research may consider using non-homothetic preferences when modeling the long-run demand for non-renewable resources.² For example, the environmental effects of non-renewable resource use would not only depend on endogenous technological change (e.g. [Acemoglu et al., 2012a](#)), but also on the level of output. Moreover, the likelihood of resource wars would not only be driven by the price elasticity of demand (as in [Acemoglu et al., 2012b](#)), but also by the elasticity of demand with regard to output. [Stefanski \(2014\)](#) introduces non-homothetic preferences in a growth model with crude oil. However, in this model non-

¹ The “intensity of use” measures how many units of a certain material are used to produce one unit of output (see [Malenbaum, 1978](#); [Tilton, 1990](#), and others).

² The literature typically models other mechanisms, namely substitution by other production factors, triggered by relative price changes (see [Solow, 1974](#); [Stiglitz, 1974](#), and the following literature) and technological change in resource use (e.g. [Acemoglu et al., 2012a](#), and others).

homothetic preferences only drive the intersectoral change in demand but not change in intrasectoral demand. Boppart (2014) provides a general growth model with non-homothetic preferences, which explains structural changes in expenditure by relative prices and income effects. The paper could serve as a valuable starting point for a growth model with non-renewable resources and non-homothetic preferences.

The paper is structured as follows. Section 2 introduces the data set. Section 3 introduces the econometric model. Section 4 presents the estimation results. Section 6 describes robustness checks, while Section 7 draws conclusions.

2. A new data set

I construct a new panel data set, which allows the regression of annual log per capita demand for a specific commodity on log per capita real manufacturing output and log real price for the period 1840 to 2010. The data set is a sample of 12 industrialized countries: Belgium, Finland, France, Germany, Italy, Japan, South Korea, the Netherlands, Spain, Sweden, the United Kingdom, and the United States. I also added three currently industrializing countries—China, India, and Brazil—whose demand for mineral commodities has substantially increased over the last two decades.

The commodity markets for aluminum, copper, lead, tin, and zinc are examined.³ These commodities were traded on the London Metal Exchange as fungible and homogeneous goods in an integrated world market over the long period considered here. They exhibit a substantial track record in industrial use. Hence, they have long-term characteristics that other mineral commodities such as iron ore, crude oil, or coal have only gained in recent times and that make a long-run analysis feasible.

The demand for a mineral commodity, the dependent variable, is derived from manufacturing sector output. The demand data capture those quantities of mineral commodities that are finished but unwrought (e.g., metal in primary shapes, such as cathodes and bars), and that manufacturers use at the first stage of production (e.g., brass mills, foundries). This is also the stage at which mineral commodities are usually traded, and it is the usual data set employed for measuring the use of mineral commodities (Tilton, 1990; U.S. Geological Survey, 2011). From 1840 to 1918, I compute the apparent usage of the respective mineral commodities, drawing on production data, as well as on import and export data from several sources. Stocks are not included in the computation of usage, due to a lack of data. As this latter measurement error is rather stochastic in nature, coefficients might be underestimated to a certain extent. From the end of World War I to today, I employ data from the German Federal Institute for Geosciences and Natural Resources (BGR, 2012). The data are mainly based on direct surveys of manufacturing industries. It has been rounded by the German Federal Institute for Geosciences and Natural Resources, which might slightly enlarge standard deviations.

I use per capita real value added in the manufacturing sector as an explanatory variable. To obtain a comparable measure across countries, I first compute the share of manufacturing in GDP from national account data from several sources. I then multiply these percentage shares with GDP data in constant International Geary-Khamis Dollars⁴ from the Maddison (2010) data set, which is a standard data set in the economic history literature. All historical national account data are based on later reconstructions and measurement errors are a potential problem. To the extent that measurement errors are stochastic, estimates will be biased towards zero and underestimate the true value. There might also be systematic measurement errors, whose biases are hard to judge.

To control for the effect of price on demand, I assemble and construct historical real prices for each country. I collect price data in local currencies for the U.S., U.K., and Germany from several sources; there are no price data series available for the other countries. However, there is strong evidence that the five mineral commodities were traded in integrated world markets with a world market price set in London over the examined time period, 1840 to 2010 (see Klovland, 2005; O'Rourke and Williamson, 1994; Labys, 2008; Stuermer and von Hagen, 2012).

I derive proxies for the national prices of the other countries by using historical exchange rates from Bordo (2001), Officer (2006, 2011), Denzel (2010), and others. I compute real prices for each country by using producer price indices from Mitchell (2003a,b, 1998), and other sources. Adjusting for inflation accounts for the erosion of the value of the respective currency, making it possible to compare price data over time. Finally, my approach neglects some price differentials due to transport costs. These appear at the price level and decrease gradually over the time period but are considered as not being substantial for these base metals in the above mentioned literature.

3. Empirical strategy

My estimation strategy builds on an extension of the partial adjustment model, which is the standard approach in empirical energy demand analysis (see Adeyemi and Hunt, 2007; Pesaran et al., 1998, 1999). I set up an auto-regressive distributed lag model (ARDL)(p, q, r) of a log linear demand function, where p, q , and r notify the number of lags included of the three explanatory variables:

³ See Table 1 for summary statistics, and the online appendix for detailed data sources and descriptions.

⁴ The International Geary-Khamis Dollar is a hypothetical unit of currency that allows for international comparison of national accounts across countries and time periods. It relies on purchasing power parity converters and is deflated with the base year 1990.

$$c_{i,t} = \sum_{j=1}^p \lambda_{ij} c_{i,t-j} + \sum_{l=0}^q \delta_{il} y_{i,t-l} + \sum_{m=0}^r \gamma_{im} p_{i,t-m} + \mu_i + \epsilon_{it}. \quad (1)$$

I explain the demand for mineral commodities $c_{i,t}$ (measured in metric tons per capita) of country i at time t by real per capita value added in the manufacturing sector $y_{i,t}$, by the real price of the respective mineral commodity $p_{i,t}$, and by its own lagged values. To capture proportional effects, I employ natural logs to all variables. I control for the effect of population growth by using per capita manufacturing output, as well as per capita demand of each mineral commodity. Country fixed effects, μ_i , capture omitted country-specific variables that are time independent. For example, a strong domestic copper mining industry might cause a generally higher level of copper demand in a country as downstream manufacturing specializes in processing copper.

Following Pesaran et al. (1998), I add a common linear time trend in the second specification and time fixed effects in the third specification. This allows me to take advantage of the panel structure of the data in a stepwise manner, and to control for a variety of omitted common factors such as technological change and spillover effects (Pesaran et al., 1998). Time fixed effects might also account for a variety of other factors such as the impact of the two world wars on the demand for mineral commodities. As they might also capture changes in world prices, they only leave those price changes in the regression that are due to changes in inflation and exchange rates. If market participants assume that these nominal shocks exhibit no long-term impact on real prices, the estimated price elasticity will be small and/or statistically insignificant in specification 3.

Reparametrizing Eq. (1), I obtain the error correction form of specification 1

$$\Delta c_{i,t} = \Phi_i(c_{i,t-1} - \theta_{0,i} - \theta_{1,i}y_{i,t} - \theta_{2,i}p_{i,t}) + \sum_{j=1}^{p-1} \lambda_{ij}^* \Delta c_{i,t-j} + \sum_{l=0}^{q-1} \delta_{il}^* \Delta y_{i,t-l} + \sum_{m=0}^{r-1} \gamma_{im}^* \Delta p_{i,t-m} + \epsilon_{it}, \quad (2)$$

where the vector θ_i captures the long-run relationship between the variables. $\theta_{1,i}$ is the long-run elasticity of demand with respect to value added by the manufacturing sector and $\theta_{2,i}$ represents the long-run elasticity of demand with respect to real price. Φ_i denotes the speed of adjustment towards the long-run equilibrium.

I introduce commonality across countries in a stepwise manner. I first employ the mean group (MG) estimator proposed by Pesaran and Smith (1995). I derive the full panel estimates of θ , Φ , δ , and γ by averaging the individual country coefficients. This estimator imposes no homogeneity restrictions on long-run or short-run restrictions. I then employ the pooled mean group (PMG) estimator proposed by Pesaran et al. (1999). The PMG estimator allows the short-run effects to vary across countries, whereas it imposes homogeneity of the coefficients for the long-run effects. For example, different economic structures across countries may affect the strength and speed at which manufacturing output and price affect the demand for mineral commodities in the short-run. The standard Hausman (1978) test is used, as proposed by Pesaran et al. (1999), to examine if long-run elasticity is, in fact, equal across the countries. If the null hypothesis of equality is not rejected, the PMG estimator is superior to the MG estimator, as it is both consistent and efficient in this case while the MG estimator is only consistent.

Using the pooled mean group (PMG) estimator, the estimated equation in specification 1 becomes

$$\Delta c_{i,t} = \Phi(c_{t-1} - \theta_0 - \theta_1 y_t - \theta_2 p_t) + \sum_{j=1}^{p-1} \lambda_{ij}^* \Delta c_{i,t-j} + \sum_{l=0}^{q-1} \delta_{il}^* \Delta y_{i,t-l} + \sum_{m=0}^{r-1} \gamma_{im}^* \Delta p_{i,t-m} + \epsilon_{it}. \quad (3)$$

Finally, I also run the standard dynamic fixed effects estimator, which restricts the long-run and short-run coefficients as well as the adjustment coefficient, making them equal across the range of countries.

There is the well-known identification problem in estimating energy demand elasticities. There might be reverse causality running from the demand variable to the price variable. The demand curve will only be identified if national prices closely follow international prices or supply is highly elastic (Pesaran et al., 1998). In this study, domestic prices follow—partly by construction—international prices as these markets have been fairly well-integrated at the global level (see Klovland, 2005; O'Rourke and Williamson, 1994; Labys, 2008; Stuermer and von Hagen, 2012) so that national demand does not affect national prices. In addition, the supply of mineral commodities is highly elastic in the long-run, according to Radetzki (2008); Krautkraemer (1998) and others (see also the theoretical argument in Stuermer and Schwerhoff (2012)), and I believe it is therefore plausible to assume that demand from a single country does not cause a long-term change in world market price.

My model might omit important variables with the potential to cause endogeneity issues and bias the estimation results. For example, to derive consumption data for the 19th century, I have computed consumption from production, export and import data for a majority of countries. Hence, inventory changes are included in the consumption data, as there is no separate information on inventories. There might therefore be reverse causality running from the demand variable to the price variable, because the demand variable also includes inventory changes. Changes in expectations about future supply and demand, changes in real interest rates, or “speculative” behavior of market participants might also affect both consumption and prices at the same time (see also Kilian and Murphy, 2014). As Pesaran et al. (1999) points out, omitted variables or measurement errors correlated with the regressors might lead to biased estimates if the bias-inducing correlations are systematic across different groups in the sample. Unfortunately, there is no way to test for this. I address this issue by running a split sample analysis for the period until 1939, which might include inventories in the consumption data, and for the following

period until the end of the sample, where data are largely based on manufacturing surveys and should not include inventories.

I use unbalanced panel data for each of the five mineral commodities. The time dimension is relatively large, while the cross-sectional dimension is not large with the number of countries $N = 15$. The incidental parameter problem (Nickell, 1981), which affects dynamic panel data models with small T and large N , is, therefore, not an issue. The common long-run coefficients of θ_i from the PMG estimator are consistent as long as $T \rightarrow \infty$, even if N is small (Pesaran et al., 1999). The ARDL specification makes unit root pretesting of the variables unnecessary. Pesaran and Smith (1995) and Pesaran (1997) show that the method is valid whether or not the variables follow a unit root process. This is based on the assumptions that there is in fact a long-run relationship, that regressors are strictly exogenous, and that there is no serial correlation in the residuals. The existence of a long-run relationship requires the adjustment coefficient to fulfill $-2 < \Phi_i < 0$ (Loayza and Ranci  re, 2006).

I model time fixed effects by expressing all variables as deviations from their respective cross-sectional means in each period in line with Pesaran et al. (1999). Such a procedure reduces common time specific effects and makes PMG estimates consistent. PMG estimation assumes that regression residuals are independent across countries. Non-zero error covariances may arise due to the omission of these common effects (Pesaran et al., 1999). The disadvantage of including time fixed effects is that they also control for changes in the world market price, leaving only those price changes in the regression caused by changes in inflation and exchange rates. If market participants assume that these nominal shocks exhibit no long-term impact on prices, the estimated price elasticities will be small and/or statistically insignificant.

Determining the lag order by information criteria on a country-by-country basis reveals significant differences across countries. However, to make regression results for the short-run and long-run parameters comparable, a common lag structure is imposed across countries. The benchmark model is an ARDL(4,4,2) model, with the inclusion of four lags of mineral commodity demand and of manufacturing output, and two lags of mineral commodity prices, respectively, in Eq. (3). I use a comparatively long lag structure to allow for rich dynamics and to account for possible serial correlation in the data. I check the robustness of my results with respect to a different choice of lag lengths as I present estimation results for ARDL(1,1,1) and ARDL(3,3,3) specifications of Eq. (3).

4. Estimation results

I find pronounced differences in the estimated long-run elasticities of demand with regard to manufacturing output across the five examined mineral commodities. Table 2 shows the estimation results for pooled mean group estimations of the first specification.⁵ I find a rather high coefficient in the case of aluminum: A 1.0 percent increase in manufacturing output leads to a more than 1.5 percent increase in aluminum demand. The estimate for copper yields a long-run elasticity of demand with regard to manufacturing output close to unity. The estimated long-run manufacturing output elasticities of lead, tin, and zinc are far below, 0.4, 0.6, and 0.7, respectively.

These results imply that the demand for aluminum increases at a higher rate than manufacturing output and, hence, the manufacturing sector's demand for aluminum increases over the course of industrialization. The demand for copper increases approximately at the same rate as manufacturing output. Thus, the copper intensity of the manufacturing sector is relatively constant. The demand for lead, tin, and zinc increases at a lower rate than manufacturing output. The intensity of use of these commodities declines over the course of industrialization.

The estimated elasticities are generally in line with the evidence on the use of the commodities. Aluminum is mainly used for the production of high technology goods such as airplanes, electronics, or machinery. It seems plausible that the relative demand for these goods increases over the course of industrialization, as consumer preferences shift to higher-tech products. Copper is very versatile, and the manufacturing sector employs it in the production of a broad array of products in electronics, construction, and transportation as well as in machinery (see Radetzki, 2009; Mardones et al., 1985). This explains why the overall demand for products that incorporate copper stayed relatively constant. The main application of zinc is in galvanization. Its use is strongly linked to products of the steel industry that lose importance as preferences shift over the course of industrialization (see Gupta, 1982; Jolly, 1997). Finally, it seems reasonable that products which extensively use tin and lead have a relatively strong demand at the early stages of industrialization but then lose relative importance. Lead is used in the production of a variety of manufactured goods such as pipes and batteries. It is an important alloy, especially in solder, but has hazard properties and in more developed, high value-added situations it falls out of favor, in part because of subsequent disposal, which is costly (Krebs, 2006). Tin is mainly used in the packaging industry as tinplate, a tin coating on thin steel. It is also employed as an alloy with lead as solder and is applied in different alloys, of which bronze is the most important (Krebs, 2006; Stuermer and von Hagen, 2012). The results suggest that the decreasing use of lead in gasoline, paint pigments, and pipes due to negative health and environmental impacts is substantially driven by a higher per capita manufacturing output and the related change in consumer preferences.

⁵ To ease exposition, I refrain from presenting the regression results for all three estimators in this chapter. The results for the mean group estimations and the dynamic fixed effects estimations can be found in the section on robustness checks and in the online appendix. Please note that I report the Hausman test to examine if long-run elasticity is, in fact, equal across the countries. If the null hypothesis of equality is not rejected, the PMG estimator is superior to the MG estimator as it is both consistent and efficient in this case, while the MG estimator is only consistent.

Table 1
Summary statistics.

Variable	Mean	Std. dev.	Min	Max	N
Per capita value added by manufacturing (GK-\$)	1694	1269	83	6566	1561
Per capita use of aluminum (mt/person)	.0059	.0075	.0000	.0490	1300
Per capita use of copper (mt/person)	.0049	.0061	.0000	.0402	1632
Per capita use of lead (mt/person)	.0027	.0020	.0000	.0079	1437
Per capita use of tin (mt/person)	.0002	.0001	.0000	.0011	1562
Per capita use of zink (mt/person)	.0033	.0043	.0000	.0384	1621
Real price of aluminum (local currencies per mt)	980.7	5046	.7743	140,411	1442
Real price of copper (local currencies per mt)	7793	32,979	0.92	388,292	1535
Real price of lead (local currencies per mt)	2404	9984	.2790	129,483	1530
Real price of tin (local currencies per mt)	26,686	109,341	2.53	1,068,194	1520
Real price of zinc (local currencies per mt)	244	503	.4708	3798	1518

Please note that these variables enter the regression in logs.

Table 2
Estimates of the long-run manufacturing output and price elasticities for Specification 1.

	Aluminum	Copper	Lead	Tin	Zinc
Country fixed effects	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	No	No	No	No
Manufacturing (log) (θ_1)	1.495*** (0.087)	0.933*** (0.057)	0.428*** (0.058)	0.580*** (0.044)	0.745*** (0.033)
Real price (log) (θ_2)	−0.691*** (0.177)	−0.403*** (0.082)	−0.227** (0.093)	−0.104 (0.063)	−0.035 (0.083)
Adjustment coefficient (Φ)	−0.106*** (0.021)	−0.130*** (0.033)	−0.079*** (0.020)	−0.147*** (0.045)	−0.104*** (0.032)
Constant (θ_0)	0.038 (0.062)	0.171** (0.072)	0.051** (0.022)	0.127*** (0.048)	0.036 (0.024)
Observations	1091	1322	1178	1260	1334
Joint Hausman test-stat.	2.072	3.256	0.960	0.960	1.283
p-value	0.355	0.196	0.619	0.619	0.527
log likelihood	521.5	577.1	540.3	396.0	690.2

Notes: The table shows results from the pooled mean group (PMG) estimations of the preferred ARDL(4,4,2) model. Standard errors in parentheses. * $p < 0.1$.

** $p < 0.05$.

*** $p < 0.01$.

The estimated long-run elasticities of demand with regard to real prices are inelastic for all examined mineral commodities. The estimated long-run price elasticity of aluminum demand is statistically highly significant and about -0.7 . This is in line with the fact that aluminum has substituted for many different materials such as composites, glass, paper, plastics, copper, and steel in a wide range of applications in manufacturing production over time (Radetzki, 2008; Chandler, 1990).

The estimated long-run price elasticity of demand of copper is lower than that of aluminum, with a statistically highly significant point estimate of about -0.4 in the first specification. This shows that copper is only moderately substitutable in its major applications. Although aluminum, plastics, and fiber optics have been substitutes for copper, especially in building materials and data transmission, its substitutability is very low in applications as a conductor of electricity (see Krebs, 2006). The estimates for the price elasticities of lead, tin and zinc demand are far lower than those for aluminum and copper, and only statistically significant at the five-percent level in the case of lead.

I find evidence for the existence of long-run relationships in all regressions, as the adjustment coefficients are all highly significant. They suggest that the estimated speed of demand adjustment is rather slow for all commodities. The estimated coefficients suggest a speed of convergence to equilibrium of about 8–15 percent per year across the five commodities. This implies that it takes about 7–13 years to revert back to equilibrium. Tin and copper markets adjust fastest, while the lead market takes the longest time to return to equilibrium. This is reasonable, given that adjustments in manufacturing capital are fairly slow and that inventories play an important role in these markets.

Hausman tests provide evidence that the long-run elasticity is, in fact, equal across the countries for all five examined commodities. As the null hypothesis of equality is not rejected, the PMG estimator is superior to the MG estimator as it is both consistent and efficient in this case, while the MG estimator is only consistent.

In the second specification, I introduce a linear time trend to account for technological change, which drives the use of these metals in all countries at the same time. This might include technological change in resource efficiency or the invention of new products. Introducing a linear time trend does not affect the pronounced differences in the estimated long-run manufacturing output elasticities of demand, which I find across the five examined mineral commodities, as Table 3 shows. Like in specification 1, aluminum has the highest estimated long-run manufacturing output elasticity of demand, while lead and tin have the lowest.

Table 3

Estimates of the long-run manufacturing output and price elasticities for Specification 2.

	Aluminum	Copper	Lead	Tin	Zinc
Country fixed effects	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	No	No	No	No
Manufacturing (log) (θ_1)	1.655*** (0.162)	1.102*** (0.134)	0.670*** (0.112)	0.595*** (0.086)	0.923*** (0.091)
Real price (log) (θ_2)	−0.773*** (0.195)	−0.440*** (0.078)	−0.213*** (0.079)	−0.112 (0.076)	−0.046 (0.078)
Linear trend	−0.008 (0.006)	−0.004 (0.004)	−0.005*** (0.002)	−0.001 (0.003)	−0.003** (0.002)
Adjustment coefficient (Φ)	−0.104*** (0.022)	−0.129*** (0.035)	−0.101*** (0.024)	−0.147*** (0.044)	−0.109*** (0.035)
Constant (θ_0)	0.998*** (0.313)	0.806** (0.366)	0.348*** (0.095)	0.252*** (0.086)	0.290* (0.152)
Observations	1091	1322	1178	1260	1334
Joint Hausman test-stat.	14.76	−100.4	6.897	2.073	10.12
p-value	0.002	1	0.075	0.557	0.018
log likelihood	522.0	577.3	542.2	396.0	691.7

Notes: The table shows results from the pooled mean group (PMG) estimations of the preferred ARDL(4,4,2) model. Standard errors in parentheses.

* $p < 0.1$.

** $p < 0.05$.

*** $p < 0.01$.

The estimated coefficients for the linear time trends are all negative and drive down the demand for aluminum by about 1 percent per year and the demand for the other commodities by less than 0.5 percent per year. However, only the ones for lead and zinc are statistically significant. This suggests that the decreasing use of lead in gasoline, paint pigments, and pipes due to negative health and environmental impacts is strongly driven by a time-related common linear trend, as various governments' regulations took effect, especially since the 1960s and 1970s (see [Smith, 1999](#)). The evidence of a linear negative trend for the usage of zinc is in line with the shrinking use of brass for all sorts of home appliances and substitution with other metals in steel production.

Hausman tests only provide evidence for homogeneity in the long-run coefficients for copper, lead, and tin demand, in this specification.

Including a linear time trend does not substantially change the estimated adjustment coefficients. They are approximately as low as in the first specification, annual rates of roughly 10 percent, which imply that it takes about 10 years for these commodity markets to revert back to equilibrium (see [Table 3](#)). Again, adjustment is fastest in the tin and copper market and slowest in the lead market.

In the third specification, time fixed effects are employed instead of linear time trends to control for common shocks from technological change. They also account for other factors which might affect the demand for these commodities in all countries at the same time, e.g., the two world war periods. The general picture stays the same (see [Table 4](#)). There are pronounced differences in the estimated long-run manufacturing output elasticities of demand across the five examined mineral commodities. As before, aluminum has a high estimated long-run manufacturing output elasticity of demand of about 1.2. Copper's manufacturing output elasticity of demand is unchanged at 0.9. Lead is slightly above the benchmark results, while zinc is slightly lower. Tin's elasticity of demand becomes significantly higher than in the other two specifications, at about unity. The International Tin Organization has long affected the tin market, which might be captured by the time fixed effects (see [Stuermer, forthcoming](#)).

As in the specifications before, the estimates of price elasticities are very low. While the price elasticity of aluminum demand is about −0.5, those for the other metals are around or below −0.2. They have all come down a bit compared with the previous specifications. The estimated adjustment coefficients still suggest that it takes a long time before demand adjusts to equilibrium (see [Table 4](#)). Hausman tests only provide evidence for homogeneity in the long-run coefficients for aluminum, lead, and tin demand, in this specification.

5. Split sample analysis

Splitting the sample into pre- and post-World War II periods shows important structural changes in the dynamics, while the overall picture stays the same (see [Tables 5 and 6](#)).⁶ Please note that only Germany, France, the U.K., Italy, Japan, and the U.S. are included in the pre-World War II sample due to data constraints.

During the pre-World War II period, the elasticities of demand with respect to manufacturing output are slightly higher for copper, tin, and zinc, than during the full sample period and the post-World War II period. This lines up well with the notion that manufacturing output has become less resource intensive and more capital and human capital intensive in

⁶ Please see the online appendix for results for the other specifications.

Table 4

Estimates of the long-run manufacturing output and price elasticities for Specification 3.

	Aluminum	Copper	Lead	Tin	Zinc
Country fixed effects	Yes	Yes	Yes	Yes	Yes
Time fixed effects	Yes	Yes	Yes	Yes	Yes
Manufacturing (log) (θ_1)	1.184*** (0.077)	0.910*** (0.087)	0.658*** (0.126)	0.997*** (0.068)	0.595*** (0.123)
Real price (log) (θ_2)	−0.531*** (0.129)	−0.070** (0.030)	−0.140*** (0.050)	−0.225*** (0.033)	0.031 (0.052)
Adjustment coefficient (Φ)	−0.115*** (0.036)	−0.133*** (0.051)	−0.110*** (0.025)	−0.146*** (0.030)	−0.078*** (0.026)
Constant (θ_0)	0.067 (0.053)	0.010 (0.039)	0.005 (0.016)	0.049 (0.033)	−0.005 (0.033)
Observations	1091	1322	1178	1260	1334
Joint Hausman test-stat.	0.436	10.47	0.746	0.363	7.953
p-value	0.804	0.005	0.689	0.834	0.019
log likelihood	593.4	590.5	505.5	432.4	797.3

Notes: The table shows results from the pooled mean group (PMG) estimations of the preferred ARDL(4,4,2) model. Standard errors in parentheses. * $p < 0.1$.** $p < 0.05$.*** $p < 0.01$.**Table 5**

Estimates of the long-run manufacturing output and price elasticities for Specification 1 during the pre-World War II period.

	Aluminum	Copper	Lead	Tin	Zinc
Country fixed effects	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	No	No	No	No
Manufacturing (log) (θ_1)	0.375 (3.688)	1.833*** (0.056)	0.295*** (0.109)	0.764*** (0.035)	1.110*** (0.047)
Real price (log) (θ_2)	6.020 (6.508)	0.345*** (0.094)	−0.114 (0.173)	−0.237*** (0.042)	0.184*** (0.063)
Adjustment coefficient (Φ)	−0.028* (0.016)	−0.314** (0.130)	−0.256* (0.131)	−0.389* (0.219)	−0.374*** (0.142)
Constant (θ_0)	−0.792 (0.758)	−2.521*** (0.939)	0.326* (0.181)	0.523** (0.258)	−1.065*** (0.365)
Observations	208	356	197	343	345
Joint Hausman test-stat.	0.287	0.312	2.701	5.787	1.418
p-value	0.866	0.856	0.259	0.055	0.492
log likelihood	55.32	110.3	148.1	165.8	157.7

Notes: The table shows results from the pooled mean group (PMG) estimations of the preferred ARDL(4,4,2) model. Standard errors in parentheses.

* $p < 0.1$.** $p < 0.05$.*** $p < 0.01$.**Table 6**

Estimates of the long-run manufacturing output and price elasticities for Specification 1 during the post-World War II period.

	Aluminum	Copper	Lead	Tin	Zinc
Country fixed effects	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	No	No	No	No
Manufacturing (log) (θ_1)	1.372*** (0.099)	0.758*** (0.095)	0.568*** (0.043)	0.589*** (0.038)	0.873*** (0.075)
Real price (log) (θ_2)	−0.160 (0.106)	−0.399*** (0.090)	−0.142*** (0.033)	−0.213*** (0.019)	−0.278*** (0.090)
Adjustment coefficient (Φ)	−0.149*** (0.027)	−0.112*** (0.031)	−0.127** (0.051)	−0.221*** (0.063)	−0.108*** (0.031)
Constant (θ_0)	−0.152*** (0.063)	0.294*** (0.084)	−0.082** (0.036)	0.340*** (0.106)	0.033 (0.025)
Observations	826	827	827	763	826
Joint Hausman test-stat.	1.366	0.824	0.945	0.481	1.868
p-value	0.505	0.662	0.623	0.786	0.393
log likelihood	714.3	739.5	693.5	470.8	821.5

Notes: The table shows results from the pooled mean group (PMG) estimations of the preferred ARDL(4,4,2) model. Standard errors in parentheses. * $p < 0.1$.** $p < 0.05$.*** $p < 0.01$.

the second half of the 20th century. At the same time, the manufacturing output elasticity of demand for aluminum is much smaller. This seems plausible, as aluminum was rarely used in manufacturing products before World War II. The coefficients for the post-World War II period are roughly in line with the full sample benchmark results.

Price elasticities for the pre- and post-World War II periods are basically in line with the full sample benchmark results. Only the price elasticity of aluminum is much higher during the pre-World War II period and much lower in the post-World War II period, but statistically insignificant in both cases. This might point to the important role of strongly falling prices in stimulating aluminum demand during the earlier period. An implausible result is the positive price elasticity for copper in the pre-World War II period. It is important to note that there are only a small number of observations in this sample and possibly problems with outliers.

Hausman tests do not reject the hypothesis that there is homogeneity across countries for all commodities during the pre-World War II and the post-World War II period.

Finally, the speed of adjustment is notably faster during the earlier period of the sample than during the post-World War II period. One explanation is that the capital intensity of the sector was lower, making adjustment of capital during the pre-World War II period easier than during the most recent period.

6. Sensitivity analysis

I check the robustness of my results with respect to the use of other estimators and different choices of lag length.

Table 7 compares the results of the pooled mean group estimator (PMG) to those of the mean group (MG) estimator and the dynamic fixed effects (DFE) estimator for the case of copper. The results for the other commodities are in the online appendix.⁷ The MG estimator does not impose any homogeneity across slopes and error variances, whereas the DFE estimator assumes homogeneity across all slopes and error variances.

The estimated long-run manufacturing output and price elasticities of demand are relatively robust across the different estimators. As expected, the standard errors of the MG estimates are larger and the coefficients are not often statistically significant. Pooling sharpens the estimates considerably as they are more robust to outliers. In the case of aluminum, the effect of the outlier Belgium is obvious and distorts the estimates. The estimated coefficients for the speed of adjustment are in all cases fairly low but significant.

Joint Hausman tests (see tables in the online appendix) do not reject the hypothesis of homogeneity of all long-run coefficients at conventional levels of significance, when the PMG estimates are compared to the MG estimates. As PMG estimates are more efficient than MG estimates, they ought to be preferred. Overall, the joint Hausman tests provide evidence that the data is mostly not violated by relying on PMG rather than MG estimates for all mineral commodities in the baseline regressions and in those with time fixed effects (see also Pesaran et al., 1999). The results for the regressions with linear trends are mixed.

The model is re-estimated using ARDL(1,1,1) and ARDL(3,3,3) configurations (see Tables 8 and 9) for the case of copper and the respective tables in the online appendix for the other commodities studied. Smaller lag lengths yield qualitatively similar results for all mineral commodities except tin, where the price elasticity becomes insignificant in the case of ARDL(3,3,3). Again, the null hypothesis of the Hausman test is mostly not rejected in the baseline specification and in the specification with time fixed effects. However, it is rejected in some specifications including linear trends.

The adjustment coefficients are statistically significant in all estimations, showing strong evidence for long-run relationships between variables.

7. Conclusion

This paper provided empirical evidence of the effect of industrialization on the derived demand for mineral commodities in a new data set that begins in 1840. After controlling for sectoral shifts, substitution, real prices, and technological change, I find substantial heterogeneity in the estimated long-run effects of manufacturing output on the demand across commodities. A 1 percent increase in per capita manufacturing output leads to a 1.5 percent increase in aluminum demand and a roughly 1 percent rise in copper demand. Estimated elasticities for lead, tin, and zinc are below unity. Common linear trends in resource efficiency and the invention of new products, which are time dependent, affect demand for zinc and lead.

My results imply substantially different paths of consumption of mineral commodities by emerging economies such as China. A slowdown in the growth rate of Chinese manufacturing output will have a stronger negative effect on the demand for aluminum or copper than for zinc, tin, or lead. At the same time, total consumption of lead, tin, and zinc will grow at a lower rate in the long run than copper and aluminum, because lead, tin and zinc have a decreasing material intensity of use over the course of industrialization. This shows that commodity exporting countries will be affected differently by a slowdown in China's industrialization process, depending on their mix of commodity exports.

⁷ Please find the online appendix at <https://sites.google.com/site/mstuermer1/research-1>.

Table 7

Estimation results from the mean group (MG), pooled mean group (PMG), and dynamic fixed effects (DFE) estimators of the preferred ARDL(4,4,2) model for copper.

Variables	MG	PMG	DFE	MG	PMG	DFE	MG	PMG	DFE
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	No	No	No	No	No	Yes	Yes	Yes
Manufacturing (log) (θ_1)	1.100*** (0.162)	0.933*** (0.057)	1.086*** (0.083)	1.065*** (0.158)	1.102*** (0.134)	1.071*** (0.171)	1.379*** (0.227)	0.910*** (0.087)	1.126*** (0.185)
Real copper price (log) (θ_2)	-0.161 (0.154)	-0.403*** (0.082)	-0.164 (0.166)	-0.173* (0.098)	-0.440*** (0.078)	-0.159 (0.172)	-0.290** (0.130)	-0.070** (0.030)	0.005 (0.079)
Linear trend				0.007* (0.004)	-0.004 (0.004)	0.000 (0.004)			
Adjustment coefficient (Φ)	-0.251*** (0.046)	-0.130*** (0.033)	-0.102*** (0.014)	-0.304*** (0.047)	-0.129*** (0.035)	-0.102*** (0.015)	-0.278*** (0.066)	-0.133*** (0.051)	-0.095*** (0.015)
Constant (θ_0)	-0.413 (0.398)	0.171** (0.072)	-0.142 (0.116)	-6.777* (3.599)	0.806* (0.366)	-0.178 (0.370)	0.287* (0.168)	0.010 (0.039)	-0.000 (0.006)
Observations	1322	1322	1322	1322	1322	1322	1322	1322	1322
Joint Hausman test-stat.		3.256			-100.4			10.47	
p-value		0.196			1			0.005	
log likelihood		577.1			577.3			590.5	

Standard errors in parentheses.

* $p < 0.1$.

** $p < 0.05$.

*** $p < 0.01$.

Table 8

Estimated long-run manufacturing output and price elasticities of copper demand in the ARDL(1,1,1) model using the mean group (MG), the pooled mean group (PMG), and the dynamic fixed effects (DFE) estimators.

Variables	MG	PMG	DFE	MG	PMG	DFE	MG	PMG	DFE
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	No	No	No	No	No	Yes	Yes	Yes
Manufacturing (log) (θ_1)	1.275*** (0.228)	1.047*** (0.038)	1.093*** (0.092)	1.365*** (0.200)	0.955*** (0.093)	1.079*** (0.200)	1.609** (0.714)	1.045*** (0.083)	1.158*** (0.218)
Real copper price (log)(θ_2)	-0.372** (0.157)	-0.249*** (0.069)	-0.229** (0.113)	-0.341** (0.156)	-0.233*** (0.069)	-0.225** (0.112)	-0.195* (0.107)	-0.030 (0.025)	0.018 (0.138)
Constant (θ_0)	-0.373 (0.357)	-0.059 (0.041)	-0.146 (0.123)	-7.117 (7.159)	-0.456*** (0.113)	-0.190 (0.824)	0.156 (0.096)	0.001 (0.025)	0.002 (0.002)
Linear Trend				0.005 (0.010)	0.002 (0.002)	0.000 (0.006)			
Adjustment coefficient (Φ)	-0.267*** (0.034)	-0.160*** (0.027)	-0.132*** (0.024)	-0.306*** (0.033)	-0.161*** (0.028)	-0.132*** (0.025)	-0.236*** (0.040)	-0.165*** (0.035)	-0.125*** (0.024)
Observations	1399	1399	1399	1399	1399	1399	1399	1399	
Joint Hausman test-stat.		1.076			27.50			4.080	
p-value		0.584			4.61e-06			0.130	
log likelihood		414.8			415.2			469.8	

Standard errors in parentheses.

* $p < 0.1$.

** $p < 0.05$.

*** $p < 0.01$.

The observed heterogeneity implies large differences in the amplitude of demand shocks across the different examined commodities. This contributes to the understanding of differences in price volatility across commodities. The longitude of price fluctuations in these markets is partially explained as the estimated systems adjust to equilibrium over roughly 7–13 years. The estimated long-run price elasticities of demand are highly inelastic for the examined mineral commodities. This shows that these mineral commodities are relatively essential to manufacturing output, as the processing industry changes its use slowly in response to price.

Finally, my results have strong theoretical implications for modeling the long-run demand for mineral commodities. The heterogeneity in the effects of manufacturing output on the demand across commodities points to changes in the product composition of the manufacturing sector, which thereby variously affect the derived demand for mineral commodities during the process of industrialization. This suggests that future work on models which include mineral commodities or non-renewable resources may consider non-homothetic preferences to account for this heterogeneity.

Table 9

Estimated long-run manufacturing output and price elasticities of copper demand in the ARDL (3,3,3) model using the mean group (MG), pooled mean group (PMG), and dynamic fixed effects (DFE) estimators.

Variables	MG	PMG	DFE	MG	PMG	DFE	MG	PMG	DFE
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time fixed effects	No	No	No	No	No	No	Yes	Yes	Yes
hline Manufacturing (log) (θ_1)	1.119*** (0.153)	0.963*** (0.053)	1.068*** (0.080)	−1.013 (2.137)	1.314*** (0.107)	1.003*** (0.162)	1.240*** (0.212)	0.856*** (0.088)	1.078*** (0.173)
Real copper price (log) (θ_2)	−0.106 (0.131)	−0.220*** (0.082)	−0.143 (0.159)	1.987 (2.116)	−0.382*** (0.073)	−0.119 (0.167)	2.061 (2.275)	−0.066 (0.042)	0.009 (0.070)
Constant (θ_0)	−0.829* (0.455)	0.044 (0.045)	−0.144 (0.118)	−1.311 (5.818)	2.328** (1.007)	−0.312 (0.383)	−0.160 (0.394)	0.008 (0.039)	−0.000 (0.005)
Linear trend				−0.043 (0.050)	−0.010*** (0.003)	0.002 (0.004)			
Adjustment coefficient (Φ)	−0.252*** (0.045)	−0.153*** (0.037)	−0.107*** (0.014)	−0.249*** (0.040)	−0.154*** (0.042)	−0.108*** (0.014)	−0.232*** (0.043)	−0.164*** (0.049)	−0.099*** (0.014)
Observations	1351	1351	.	1351	1351	.	1351	1351	.
Joint Hausman test-stat.		1.633			38.83			3.907	
p-value		0.442			1.89e−08			0.142	
log likelihood		578.6			580.8			622.8	
Observations	1213	1213	1213	1213	1213	1213	1213	1213	1213
Joint Hausman test-stat.		2.827			20.72			0.125	
p-value		0.243			0.000,120			0.939	
log likelihood		481.2			485.1			427.5	

Standard errors in parentheses.

* $p < 0.1$.

** $p < 0.05$.

*** $p < 0.01$.

Appendix A. Supplementary material

Supplementary data associated with this article can be found, in the online version, at <http://dx.doi.org/10.1016/j.jimonfin.2017.04.006>.

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