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Estimating Cross-Industry Cross-Country Interaction Models Using
Benchmark Industry Characteristics

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Abstract

Cross-industry cross-country models are applied widely in economics. For example, to investigate the effect of financial development on economic growth or the effect of institutional quality on international trade. The literature estimates the effect of interest by examining the interaction between country characteristics—for example, financial development or institutional quality—and theoretically relevant technological industry characteristics—for example, dependence on external finance or relationship-specific inputs. As the relevant industry characteristics are unobservable for most countries, they are proxied by industry characteristics of a benchmark country. We analyze this approach when there is cross-country heterogeneity in technological industry characteristics. First, we show that the estimation approach in the literature is biased and that the bias cannot be signed if technologically similar countries are similar in terms of other characteristics. Second, we derive necessary and sufficient conditions for identification of the effect of interest. Third, we use the new identification approach to reestimate the impact of institutional quality on comparative advantage in industries that rely on relationship-specific inputs.

Keywords: Economic Growth, International Specialization and Trade, Country-Specific Technology, Financial Development, Institutions

JEL classification numbers: G30, F10, O40

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1 Introduction

Empirical work in economics has relied extensively on cross-industry cross-country models over the past 20 years. These models relate cross-country differences in industry performance—for example, industry growth or industry exports—to an interaction between (i) country characteristics like financial development, institutional quality, or human capital and (ii) theoretically relevant technological industry characteristics like external-finance dependence, reliance on certain inputs, or skill intensity. The approach has proven useful for examining a wide variety of questions, reviewed below. Two strands of research stand out. First, following Rajan and Zingales (1998), cross-industry cross-country models are used to study how financial development, property rights protection, contract enforcement, and human capital affect industry employment, output, and value added growth. Second, cross-industry cross-country models serve as the basis for empirical studies of the institutional determinants of comparative advantage; see Nunn and Trefler (2014) for a review. For example, Nunn (2007) uses the approach to show that institutional quality is a source of comparative advantage in industries that rely on relationship-specific inputs.

The theoretically relevant technological industry characteristics in the cross-industry cross-country literature depend on the economic question asked, but a common feature is that they are unavailable for most countries. This is for two main reasons. First, there is little industry data for most countries. Second, technological industry characteristics must be inferred from observed, endogenous industry characteristics (e.g., Rajan and Zingales, 1998). Such inference is challenging for countries where firms have adapted to, for example, low financial development, institutional quality, or human capital. As a result, the cross-industry cross-country literature generally proceeds by treating technological industry characteristics as unobservable for all countries except a highly-developed benchmark country with relatively undistorted markets, usually the USA. The effect of interest is then estimated using the technological industry characteristics of the benchmark country as proxies for the technological industry characteristics of all other countries.

To better understand this estimation approach, we study cross-industry cross-country models with two main features. First, following the literature, the relevant technological industry characteristics are unobservable for all but a benchmark country. Second, there is some cross-country heterogeneity in technological industry characteristics (e.g., Bernard and Jones, 1996; Acemoglu and Zilibotti, 2001; Schott, 2004; Caselli, 2005).

We show that the estimation approach used in the cross-industry cross-country literature results in a bias shaped by two countervailing forces. A key determinant of the relative strength of these forces—and hence the ultimate bias of the estimation approach—is whether technologically similar countries are similar in terms of other characteristics, for example,

their levels of financial development, institutional quality, or human capital. If there is no relationship between how similar countries are technologically and how similar they are in terms of other characteristics, the approach in the literature yields estimates of the effect of interest that are biased towards zero (attenuated). That estimates may be attenuated is generally understood in the cross-industry cross-country literature and explained in terms of a classical measurement error bias due to the industry characteristics of the benchmark country measuring the industry characteristics of other countries with some error.¹ On the other hand, if technologically similar countries are similar in terms of other characteristics, the approach in the cross-industry cross-country literature may yield amplified or spurious estimates. There appears to be no discussion of this issue in the literature.²

As the estimation approach used in the cross-industry cross-country literature does not identify the effect of interest when there is cross-country heterogeneity in technological industry characteristics, it is important to develop alternatives. To do so, we first provide an analysis of identification. Identification is substantially more challenging with cross-country heterogeneity in technological industry characteristics than without, and sometimes exact identification is impossible. Our analysis develops necessary and sufficient conditions for exact identification and provides bounds when exact identification is impossible. We use the new identification approach to reestimate the effect of institutional quality on comparative advantage in relationship-specific-input intensive industries in Nunn (2007).

The rest of the paper is structured as follows. Section 2 reviews applications of cross-industry cross-country models. Section 3 examines the estimation approach in the literature. Section 4 discusses identification of the effect of interest. Section 5 uses our identification results to reestimate Nunn (2007). Section 6 provides the conclusion.

2 Applications in the Literature

Cross-industry cross-country models have been applied extensively in many areas of economics. Our review is only meant to illustrate the range of applications. Appendix Table 1 provides brief summaries of more papers in the literature.

The Economic Effects of Financial Markets. Starting with the influential work of Rajan and Zingales (1998), who showed that financial development exerts a disproportionately

¹For example, see Rajan and Zingales (1998), p. 567. Attenuated estimates are sometimes seen as a relatively minor drawback as they can be interpreted as lower bounds on the strength of true effects.

²The literature does discuss the possibility of endogeneity bias (endogenous country characteristics or industry characteristics of the benchmark country) or omitted variable bias (relevant industry-country interactions are omitted). We abstract from these well-understood sources of estimation bias by assuming that country as well as industry characteristics of the benchmark country are exogenous and that there is a single relevant industry-country interaction.

large impact on sales growth in industries that depend more on external sources of finance, cross-industry cross-country models have been applied extensively to investigate the effects of financial markets on economic growth, firm entry and exit, investment, and innovation. For example, Fisman and Love (2003) find that in countries with less developed financial markets, industries that rely more on trade credit grow faster, and Fisman and Love (2007) show that better developed financial markets spur growth in industries facing better global growth opportunities. Claessens and Laeven (2003), Braun and Larrain (2005), and Lei, Qiu, and Wan (2018) extend the cross-industry cross-country model of Rajan and Zingales (1998) to account for the role of intangible assets. Brown, Martinson, and Petersen (2013), Hsu, Tian, and Xu (2014), and Acharya and Zu (2017) use cross-industry cross-country models to examine the impact of financial markets on innovation. Pagano and Shivardi (2003), Aghion, Fally, and Scarpetta (2007), and Beck, Demirgüç-Kunt, Laeven, and Levine (2008) analyze how financial markets affect firm entry and exit and the growth of smaller versus larger firms.

Cross-industry cross-country models are also used to examine the economic effects of specific financial market policies or institutions, such as bank recapitalizations (Laeven and Valencia, 2013), insider trading legislation (Edmans, Jayaraman, and Schneemeir, 2017), and collateral laws (Calomiris, Larrain, Liberti, and Sturgess, 2017). A more recent strand of research employs cross-industry cross-country models to assess the effects of financial crises and capital account liberalization on macroeconomic performance (Dell’Ariccia, Detragiache, and Rajan, 2008; Iacovone and Zavacka, 2009; Duchin, Ozbas, and Sensoy, 2010; Claessens, Tong, and Wei, 2012; Larrain and Stumpner, 2018).

International Specialization and Trade. Cross-industry cross-country models are widely used to examine the determinants of international trade and international specialization. Levchenko (2007) and Nunn (2007) use cross-industry cross-country models to examine the effect of institutional quality on international specialization (see also Ferguson and Formai, 2013; Nunn and Trefler, 2014). Manova (2008, 2013) uses cross-industry cross-country models to link financial development to the patterns of international trade (see also Chan and Manova, 2015; Manova, Wei, and Zhang, 2015; Claessens, Hassib, and van Horen, 2017; Crin and Ogiliri, 2017). Ciccone and Papaioannou (2009) and Debaere (2015) use cross-industry cross-country models to examine the effects of human capital and natural resources on international specialization. Cingano, Leonardi, Messina and Pica (2010), Mueller and Phillippon (2011), Cuñat and Melitz (2012), Tang (2012), Griffith and Macartney (2014), and Broner, Bustos, and Carvalho, (2016) use cross-industry cross-country models to examine role of labor-market and environmental regulation for international trade.

Other Applications. Cross-industry cross-country models have proven useful for examining a surprisingly wide variety of additional economic questions. For example, Alfaro and Charlton (2009), Carluccio and Fally (2012), Basco (2013), Blyde and Danielken (2015), Paunov (2016), and Fort (2017) use cross-industry cross-country models to analyze the determinants of outsourcing, foreign direct investment, and the fragmentation of production. Pagano and Schivardi (2003), Klapper, Laeven, and Rajan (2006), Acemoglu, Johnson, and Mitton (2009), Aizenman and Sushko (2011), Bombardini, Gallipoli, and Pupato (2012), Michelacci and Schivardi (2013), Larrain (2014), and Aghion, Howitt, and Prantl (2014) use cross-industry cross-country models to analyze the economic consequences of cross-country differences in firm size distributions, entry regulation, transaction costs, risk sharing possibilities, and skill dispersion. Rajan and Subramanian (2010) and Chauvet and Ehrhart (2018) use cross-industry cross-country models to understand the economic effects of foreign aid, while Pierce and Snyder (2017) and Levine, Lin, and Xie (2018) use them to study the legacy of slave trade. Aghion, Farhi, and Kharroubi (2015), Aghion, Hemous, and Kharroubi (2014), and Cecchetti and Kharroubi (2018) use cross-industry cross-country models to analyze the economic effects of fiscal and monetary policy over the business cycle and Avdjiev, Bruno, Koch, and Shin (2018) to analyze the economic impact of exchange rates.

3 The Standard Benchmarking Estimator

3.1 The Model

The basis of cross-industry cross-country models are theories linking industry outcomes in different countries to an interaction between country characteristics and technological industry characteristics. For example, in Rajan and Zingales (1998), the outcome variable is industry growth and the interaction is between financial development and the external-finance dependence of industries. In Nunn (2007), the outcome variable is industry exports and the interaction is between institutional quality and the intensity with which industries use relationship-specific inputs. As the main theoretical prediction concerns the effect of the interaction between country and industry characteristics, cross-industry cross-country models allow controlling for country and industry fixed effects. An empirical framework that encompasses the models in the literature is

$$y_{in} = (\alpha + \beta x_n)z_{in} + \nu_{in} \quad (1)$$

where y_{in} is the outcome in I industries indexed by i and N countries indexed by n ; x_n is the relevant country characteristic; z_{in} denotes the relevant industry characteristic in different countries; and ν_{in} captures country and industry fixed effects as well as any unobserved

determinants of industry outcomes that are independent of z_{in} . The main parameter of interest is β , the coefficient on the industry-country interaction term. The parameter α captures any direct effects of industry characteristics on outcomes. We take the relevant country characteristic x_n to be non-stochastic.

Estimation of β in (1) would be straightforward if there was data on the relevant industry characteristics z_{in} for a broad set of countries. But the necessary data is unavailable for most countries. Moreover, the cross-industry cross-country literature often focuses on technological industry characteristics that must be inferred from observed, endogenous industry characteristics. Such inference is challenging in countries where firms have adapted to, for example, low financial development, institutional quality, or human capital. As a result, the cross-industry cross-country literature generally proceeds by proxying the relevant technological industry characteristics of all countries with industry characteristics from a highly-developed benchmark country with relatively undistorted markets, usually the USA.

To better understand this benchmarking approach, it is useful to distinguish between the relevant technological industry characteristics z_{in} in (1) and observed industry characteristics \tilde{z}_{in} . Observed industry characteristics are endogenous and may therefore depend on the country characteristic x_n the model in (1) focuses on, as well as other country characteristics h_n , $\tilde{z}_{in} = g(i, x_n, h_n)$. The objective of the model in (1) is to determine the economic effects of cross-country differences in x_n through the industry-specific channel captured by z_{in} . Because \tilde{z}_{in} is endogenous to the cross-country differences in x_n , using \tilde{z}_{in} as right-hand-side industry characteristics in (1) would generally produce misleading conclusions. A better choice for the right-hand-side industry characteristics would be the hypothetical industry characteristics of countries n if they all had the same country characteristic x^* , $\tilde{z}_{in}^* = g(i, x^*, h_n)$. The industry data to infer these hypothetical industry characteristics is unavailable for most countries. As a result, the cross-industry cross-country literature generally proceeds by proxying the industry characteristics z_{in} in (1) with the industry characteristics of a highly-developed benchmark country with relatively undistorted markets.³

It is important to understand if the benchmarking approach used in the cross-industry cross-country literature can identify the effect of interest β . Clearly, the approach works if the technological industry characteristics z_{in} of all countries were identical, i.e. $z_{in} = z_i$. In this case, using the industry characteristics of a benchmark country as a proxy for the industry characteristics of all other countries would not involve any measurement error.

But as countries generally differ in a range of characteristics that could be relevant for industry structure and technology adoption, it seems extremely implausible that the relevant technological industry characteristics of all countries are identical (e.g., Bernard and Jones,

³In a few cases, z_{in} is proxied using industry data from several highly-developed countries. This does not affect our analysis below at all, except that the place of US industry characteristics would be taken by the industry characteristics of the alternative benchmark country/countries.

1996; Acemoglu and Zilibotti, 2001; Schott, 2004; Caselli, 2005).⁴ The technological industry characteristics of any benchmark country will therefore most likely be a noisy proxy for the technological industry characteristics of other countries. How good a proxy, can be expected to be country specific.

This point can be illustrated with the study of Nunn (2007). The key industry-level variable in Nunn is the technological relationship-specific-input intensity of industries and the country characteristic of interest is institutional quality. Clearly, the observed relationship-specific-input intensity of industries in a country may depend on its institutional quality, as firms might make fewer relationship-specific investments when they operate in an environment with worse institutional quality. For this reason, and because there is little industry data for most countries, Nunn proxies the technological relationship-specific-input intensity of industries of all countries by the observed relationship-specific-input intensity of US industries. However, even if all countries had the US level of institutional quality, the technological relationship-specific-input intensity of industries might still be different across countries, as industry structure and technology may not depend solely on institutional quality. Put differently, the relationship-specific-input intensity of US industries may be a noisy proxy for the technological relationship-specific-input intensity of other countries even if these countries had the institutional quality of the US. How good a proxy, depends on country characteristic other than institutional quality that affect industry structure and the technological relationship-specific-input intensity of industries.

For example, Nunn documents that industries that rely more on relationship-specific inputs also use human capital more intensively. Hence, the level of human capital of a country may affect which of the many industries with different technological relationship-specific-input intensities produce in the country. That is, the technological relationship-specific-input intensity of industries may depend on the country's human capital. As a result, the relationship-specific-input intensity of industries in a high human capital country like the US could be similar to the technological relationship-specific-input intensity of countries with similar human capital but substantially different from the technological relationship-specific-input intensity of countries with low human capital.

We want a framework that allows us to capture in a flexible way that technological industry characteristics may be more similar for some country pairs than others. The first step is to take the relevant technological industry characteristics z_{in} in (1) to be the sum of a country-specific component z_n ; a global industry-specific component z_i ; and a country-specific industry component ε_{in}

$$z_{in} = z_n + z_i + \varepsilon_{in}. \tag{2}$$

⁴In fact, the cross-industry cross-country literature has regarded this assumption as unreasonable since its beginnings, see Rajan and Zingales (1998), p. 563.

The country-specific component z_n allows us to capture all country-specific factors that shift the entire distribution of technological industry characteristics. We treat this component as non-stochastic. The global industry component z_i allows us to capture factors that make two industries i and j different from each other independently of the country where they are located.⁵ We treat this component as an independent and identically distributed random variable with $Var(z_i) > 0$. For the ε_{in} we chose a model that allows us to capture that:

- (i) How different any two industries are technologically may be country specific.
- (ii) Some countries may be more similar technologically than others.

To capture (i) and (ii) we assume that the ε_{in} in (2) are jointly normally distributed for all i and n . For any pair of countries $n \neq m$, the correlation of the ε_{in} across industries is allowed to be an arbitrary function of country characteristics

$$Corr(\varepsilon_{in}, \varepsilon_{im}) = \rho_{nm}. \quad (3)$$

As ρ_{nm} can be different for each country pair, (3) will yield a flexible model of the relationship between the characteristics of any pair of countries and their technological similarity. Across industries, the ε_{in} are taken to be independent, and

$$E(\varepsilon_{in}) = 0 \text{ and } E(\varepsilon_{in}^2) = \sigma^2. \quad (4)$$

The variance across industries of the technological industry characteristics z_{in} in (2) is $Var(z_{in}) = Var(z_i) + \sigma^2$ for all countries n . Hence, larger values of σ^2 imply that more of the heterogeneity in technological industry characteristics is country specific. The assumption that $Var(z_i) > 0$ implies that σ^2 is strictly smaller than the variance across industries of the technological industry characteristics in each country, $\sigma^2 < Var(z_{in})$ for all n . This is because $Var(z_i) > 0$ and (2) imply that at least some of the variance across industries of technological characteristics in each country reflects a global component.

If $\sigma^2 = 0$, the variance across industries of the technological industry characteristics is entirely driven by the global component in each country. As a result, there is no cross-country heterogeneity in technological differences between industries. This is because in this case, $\varepsilon_{in} = 0$ for all i, n and hence (2) implies that $z_{in} - z_{jn} = z_i - z_j$. Hence, technological differences between industries $z_{in} - z_{jn}$ do not vary at all across countries n . Because our model for z_{in} in (2) allows for a country-specific component z_n , the levels of technological industry characteristics could still vary across countries. But such cross-country heterogeneity does not play an important role in our analysis, as it is absorbed by

⁵That there must be such a global component for the estimation approach in the cross-industry cross-country literature to make sense was already pointed out in Rajan and Zingales (1998), p. 563.

the country fixed effects always present in cross-industry cross-country models.

If $\sigma^2 > 0$, there is cross-country heterogeneity in technological differences between industries. To see this, note that (2) implies $z_{in} - z_{jn} = (z_i - z_j) + (\varepsilon_{in} - \varepsilon_{jn})$ and generically $\varepsilon_{in} - \varepsilon_{jn} \neq \varepsilon_{im} - \varepsilon_{jm}$ for all $n \neq m$. To see the implications of this heterogeneity, it is useful to relate $z_{in} - z_{jn}$ for country n and any pair of industries i and j to differences in US industry characteristics, $z_{iUS} - z_{jUS}$, and differences in the global industry component, $z_i - z_j$. This yields

$$z_{in} - z_{jn} = \rho_{nUS}(z_{iUS} - z_{jUS}) + (1 - \rho_{nUS})(z_i - z_j) + u_{ijnUS} \quad (5)$$

where ρ_{nUS} refers to the correlation in (3) between the specific industry characteristics of country n and the US, and u_{ijnUS} is a random variable with $E(u_{ijnUS}) = 0$ that is independent of the z_i and the z_{iUS} .⁶ Hence, in expectation:

- (i) The difference between the technological characteristics of any two industries in country n can be thought of as a weighted average of industry differences in the US and industry differences in the global component.
- (ii) The weight on the technological industry characteristics of the US is the correlation coefficient ρ_{nUS} between the specific industry characteristics of country n and the US.

As the coefficients ρ_{nUS} in (5) can be arbitrary functions of country characteristics, our model of technological industry characteristics allows for a flexible relationship between the x -characteristics of countries and their technological similarity with the US.

It is useful to see what the model in (5) allows us to capture in the context of Rajan and Zingales (1998) and of Nunn (2007) for example. In Nunn, (5) allows us to capture that—even if all countries had the US level of institutional quality—the relationship-specific-input intensity of industries in a high human capital country like the US could be different from the relationship-specific-input intensity of industries in countries with low human capital. The estimation approach in the cross-industry cross-country literature fails to take this into account and estimates of the effect of institutional quality on industry outcomes could therefore be biased upwards or downwards. As the coefficients ρ_{nUS} in (5) can be arbitrary functions of the human capital of countries, our framework allows us to capture the relationship-specific-input intensity of industries in countries with different human capital compared to the relationship-specific-input intensity of US industries in a flexible way.⁷

In Rajan and Zingales (1998), the key industry-level variable is the external-finance intensity of industries. Rajan and Zingales point out that this variable captures technological

⁶This holds for any pair of countries n and m . It follows from (2)–(4) and joint normality of the distribution of ε_{in} for all i and n .

⁷As an aside, a country’s human capital could also affect industry outcomes through the technological human-capital intensity of industries of course.

shocks that raise an industry’s investment opportunities beyond what internal funds could support. As the benchmark country in Rajan and Zingales is the US, the external-finance intensity of industries used in their empirical analysis is that of US industries. The technological shocks affecting US industries could be similar to shocks affecting industries in countries with similar levels of, for example, human capital but quite different from shocks affecting industries in countries with low human capital. The estimation approach in the cross-industry cross-country literature does not take this into account and can therefore yield upwards or downwards biased estimates of the effect of financial development on industry outcomes. As the coefficients ρ_{nUS} in (5) can be arbitrary functions of country characteristics, our framework allows us to capture flexibly that technological shocks affecting industries in countries with low human capital could be quite different than technological shocks affecting high human capital countries like the US.

It is interesting to note that (5) does not determine whether the difference between the technological characteristics of any two industries in country n increases or decreases relative to the US as ρ_{nUS} increases. The answer depends on whether the difference between the global component of technological industry characteristics, $z_i - z_j$, is greater or smaller than the difference between the technological industry characteristics of the US, $z_{iUS} - z_{jUS}$. This gives the model additional flexibility that seems desirable. For example, consider the effect of a country’s human capital on the relationship-specific-input intensity of its industries discussed above in the context of Nunn’s (2007) study. Compared to the US, industries might be less relation-specific-input intensive in countries with low human capital. However, there seems no reason to suppose that this effect is stronger in some industries than others. Hence, the difference in the use of relation-specific inputs between industries i and j in countries with low human capital may be greater or smaller than in the US.

We could also model the endogeneity of US industry characteristics by extending our conceptual framework to include a fictional frictionless (ff) country. This would allow us to write US industry characteristics in terms of industry characteristics in the fictional frictionless country, $z_{iUS} - z_{jUS} = \rho_{USff}(z_{iff} - z_{jff}) + (1 - \rho_{USff})(z_i - z_j) + u_{ijUSff}$ where the derivation and variable definitions are analogous to (5). This results in a flexible model of how US industry characteristics—and, using (5), the characteristics of any other country—differ compared to a frictionless baseline as ρ_{USff} is allowed to be an arbitrary function of US characteristics.

3.2 Characterizing the Standard Benchmarking Estimator

We now apply the estimation approach used in the cross-industry cross-country literature to the model in (1) and (2). This yields what we refer to as the *standard benchmarking estimator*. We then discuss the forces shaping the bias of this estimator.

3.2.1 Deriving the Standard Benchmarking Estimator

The estimating equation in the cross-industry cross-country literature is

$$y_{in} = a_i + a_n + bx_n z_{iUS} + residual_{in} \quad (6)$$

where a_i and a_n are industry and country fixed effects, and z_{iUS} denotes the industry characteristics of the benchmark country (we use the subscript US as the benchmark country is usually the US). The effect of interest is captured by the coefficient b on the industry-country interaction, and the method of estimation is least squares.⁸

It is useful to write the least-squares estimator of b in (6) in terms of demeaned variables

$$\hat{b} = \frac{\frac{1}{N} \frac{1}{I} \sum_{n=1}^N \sum_{i=1}^I (z_{iUS} - \bar{z}_{US})(x_n - \bar{x})(y_{in} - \bar{y}_n - \bar{y}_i + \bar{y})}{\frac{1}{N} \frac{1}{I} \sum_{n=1}^N \sum_{i=1}^I (z_{iUS} - \bar{z}_{US})^2 (x_n - \bar{x})^2} \quad (7)$$

where \bar{y} is the average of y_{in} across industries and countries; \bar{y}_i is the cross-country average of y_{in} for industry i ; \bar{y}_n is the cross-industry average of y_{in} for country n ; \bar{z}_{US} is the cross-industry average of z_{iUS} ; and \bar{x} is the cross-country average of x_n .

To see when the standard benchmarking estimator identifies the main parameter of interest β , we consider the probability limit of \hat{b} as the number of industries goes to infinity. Substituting (1) in (7) and taking the probability limit—see the Appendix for details—yields

$$b = \text{plim}_{I \rightarrow \infty} \hat{b} = \left(1 - \frac{\sigma^2}{\sigma_{US}^2}\right) \beta + \left(\frac{\sigma^2}{\sigma_{US}^2}\right) (\alpha A + \beta B) \quad (8)$$

where σ^2 is the variance of ε_{in} and σ_{US}^2 is the variance of the US industry characteristic z_{iUS} , with $\sigma^2/\sigma_{US}^2 < 1$; α captures direct effects of industry characteristics on industry outcomes; and A and B capture the relationship between the characteristic x_n of country n and how similar the country is technologically to the US (as measured by ρ_{nUS})

$$A = \frac{\text{Cov}(x_n, \rho_{nUS})}{\text{Var}(x_n)} = \frac{\sum_{n=1}^N (x_n - \bar{x}) \rho_{nUS}}{\sum_{n=1}^N (x_n - \bar{x})^2} \quad (9)$$

⁸We assume x_n to be exogenous. In some applications in the literature, exogeneity is an issue and x_n is therefore instrumented. In these applications, our analysis applies to the reduced-form equation. We always include the US (benchmark country) as one of the countries in our analysis. The literature sometimes drops the benchmark country but, given the relatively large number of countries included, this generally makes very little difference for the estimates.

and

$$B = \frac{Cov(x_n, \rho_{nUS}x_n)}{Var(x_n)} = \frac{\sum_{n=1}^N (x_n - \bar{x})x_n\rho_{nUS}}{\sum_{n=1}^N (x_n - \bar{x})^2}. \quad (10)$$

For example, suppose that the US is a high- x country, i.e. the US has a high level of financial development, institutional quality, or human capital. Then A is positive if countries that are similar technologically to the US are also similar to the US in terms of the x -characteristic. In the typical application of cross-industry cross-country models in the literature, B would also tend to be positive in this case.⁹

An immediate implication of (8) is that the standard benchmarking estimator identifies β when there is no cross-country heterogeneity in technological industry characteristics, $\sigma^2 = 0$. In this case, the technological differences between US (benchmark country) industries are identical to the technological differences between industries of all other countries. Using US industry characteristics as a proxy for the technological industry characteristics of all other countries does therefore not involve any measurement error.¹⁰

When there is cross-country heterogeneity in technological industry characteristics, $\sigma^2 > 0$, the standard benchmarking estimator in (8) is biased and the bias is shaped by two main forces. First, how much country-specific heterogeneity there is in technological industry characteristics (captured by σ^2/σ_{US}^2). Second, how the technological similarity of countries with the US (captured by ρ_{nUS}) covaries with their characteristics x_n (captured by A and B). We now discuss the forces shaping the bias in some interesting special cases and show that the standard benchmarking estimator may be biased towards zero (attenuated); biased away from zero (amplified); or entirely spurious.

3.2.2 The Bias of the Standard Benchmarking Estimator: a First Approach

The expressions in (8)–(10) allow us to discuss the forces shaping the bias of the standard benchmarking estimator and illustrate three main types of biases.

Attenuation Bias. We start with the case that we see as corresponding to the implicit assumption in the cross-industry cross-country literature. In this case, differences between the technological industry characteristics of a country and global technological industry

⁹Theoretically, the sign of B could depend on the distribution of the x -characteristics across countries even if A is positive.

¹⁰As already mentioned, our model for z_{in} in (2) allows for a country-specific component z_n and the levels of technological industry characteristics could therefore vary across countries even if $\sigma^2 = 0$. But such cross-country heterogeneity does not play an important role in our analysis, as it is absorbed by the country fixed effects always present in cross-industry cross-country models.

characteristics are assumed to be completely idiosyncratic to the country. Put differently, the technological industry characteristics of different countries are related through global industry characteristics only, i.e. $\rho_{nm} = 0$ for all country pairs $n \neq m$.

In this case, (9) and (10) imply $A = B = 0$ and the expression for the standard benchmarking estimator in (8) simplifies to $b = \beta(1 - \sigma^2/\sigma_{US}^2)$. As already mentioned, the assumption $Var(z_i) > 0$ implies $\sigma^2/\sigma_{US}^2 < 1$ as at least some of the variation in technological industry characteristics in each country, including the US, is due to the global component. Hence, the standard benchmarking estimator b has the same sign as the parameter of interest β but is biased towards zero. This possibility is generally understood in the cross-industry cross-country literature and explained in terms of a classical measurement error bias due to US (benchmark country) industry characteristics measuring the technological industry characteristics of other countries with some error (e.g. Rajan and Zingales, 1998, p. 567). Intuitively, $\rho_{nm} = 0$ for all $n \neq m$ implies that US industry characteristics are an equally imperfect proxy for the technological industry characteristics of all other countries. US industry characteristics become a uniformly worse proxy for the technological industry characteristics of other countries for larger values of σ^2/σ_{US}^2 . As a result, the attenuation bias is stronger the greater the country-specific component of technological industry characteristics.

Spurious Interaction Effect. When there is cross-country heterogeneity in technological industry characteristics, the standard benchmarking estimator can indicate a positive effect of the country characteristic x_n on industry outcomes even though x_n does not actually enter the true model at all. To see this, suppose that $\beta = 0$, which implies that the country characteristic x_n drops out from the true model in (1). Suppose also that there is cross-country heterogeneity in technological industry characteristics, $\sigma^2 > 0$. In this case, the standard benchmarking estimator in (8) is $b = \alpha A \sigma^2 / \sigma_{US}^2$. Hence, if $\alpha A > 0$, the standard benchmarking estimator indicates a positive effect of the industry-country interaction $x_n z_{iUS}$ on industry outcomes, although the country characteristic is in fact irrelevant for industry outcomes. This is because $\alpha A > 0$ implies that cross-country heterogeneity in technology is such that industry outcomes in high- x countries are more closely correlated with US industry characteristics than industry outcomes in low- x countries.¹¹ The standard benchmarking estimator misinterprets this as a positive effect of the industry-country interaction $x_n z_{iUS}$ on industry outcomes, and therefore leads to the erroneous conclusion that the country

¹¹This could be because the technological industry characteristics of high- x countries are more similar to US industry characteristics and there is a positive direct effect of technological industry characteristics on industry outcomes ($A > 0$ and $\alpha > 0$). Alternatively, technological industry characteristics of high- x countries could be less similar to US industry characteristics and there could be a negative direct effect of technological industry characteristics on industry outcomes ($A < 0$ and $\alpha < 0$).

characteristic x_n has an effect on industry outcomes.¹²

Amplification Bias. The standard benchmarking estimator can also result in an amplification bias. To see this in the simplest case, assume there is no direct effect of industry characteristics on outcomes, $\alpha = 0$. In this case, (8) simplifies to $b = \beta [1 + (B - 1)\sigma^2/\sigma_{US}^2]$. Hence, if $B > 1$ and there is cross-country heterogeneity in technological industry characteristics ($\sigma^2 > 0$), the standard benchmarking estimator b will be an amplified version of β , $|b| > |\beta|$ and $sign(b) = sign(\beta)$.

The amplification bias of the standard benchmarking estimator is the most difficult bias to understand intuitively. At the most general level, for there to be an amplification bias, US industry characteristics must be a better proxy for the technological industry characteristics of some countries than others. Specifically, US industry characteristics must be a better proxy for the technological industry characteristics of countries that have x -characteristics similar to the US. In our framework, this is the case if countries that are more similar technologically to the US are also more similar in terms of their x -characteristics.

To see the sources of the amplification bias of the standard benchmarking estimator formally, it is useful to rewrite the model in (1) as

$$y_{in} = \gamma_n z_{in} + \nu_{in} \tag{11}$$

$$\gamma_n = \beta x_n \tag{12}$$

where we continue to assume $\alpha = 0$. We simplify further by treating the disturbance ν_{in} as an independent and identically distributed random variable. The parameters γ_n in (11) capture the effect of industry characteristics on outcomes in different countries. We refer to these parameters as country-specific slopes. The parameter β in (12) captures how these country-specific slopes covary with the country characteristic x_n .

Now imagine estimating the country-specific slopes γ_n in (11) separately for each country. As we only observe the technological industry characteristics of the US, we use US industry characteristics z_{iUS} as a proxy for the technological industry characteristics z_{in} of each country. We denote the least-squares slope estimates of γ_n by \hat{g}_n . Clearly, \hat{g}_n will generally be biased. To see the factors shaping the bias we take the probability limit of \hat{g}_n as the number

¹²More formally, when $\beta = 0$, the benchmarking estimator solely reflects the covariation between the direct effect of country-specific industry characteristics on industry outcomes $\alpha \varepsilon_{in}$ and the interaction $x_n z_{iUS}$. This covariation is $\alpha \frac{1}{N} \sum_{n=1}^N (x_n - \bar{x}) E \varepsilon_{in} (z_{iUS} - z_i) = \alpha \frac{1}{N} \sum_{i=1}^N (x_n - \bar{x}) \sigma^2 \rho_{nUS} = \alpha \sigma^2 Cov(x_n, \rho_{nUS})$ where we made use of the definition of ρ_{nUS} . Hence, as long as there is cross-country heterogeneity in technological industry characteristics, the covariation is positive if and only if $\alpha Cov(x_n, \rho_{nUS}) > 0$. Using the definition of A , this is equivalent to $\alpha A > 0$.

of industries I goes to infinity. This yields

$$g_n = \text{plim}_{I \rightarrow \infty} \hat{g}_n = \gamma_n \left[\left(1 - \frac{\sigma^2}{\sigma_{US}^2} \right) + \left(\frac{\sigma^2}{\sigma_{US}^2} \right) \rho_{nUS} \right] \quad (13)$$

where $\sigma^2/\sigma_{US}^2 < 1$. The term in square brackets turns out to be the correlation coefficient between the technological industry characteristics of country n and the US, $\text{corr}(z_{in}, z_{iUS})$. Hence, the bias of the least-squares slopes, $g_n - \gamma_n$, reflects the technological similarity between country n and the US as captured by $\text{corr}(z_{in}, z_{iUS})$. This yields two insights: (i) the more similar a country is technologically to the US (the closer $\text{corr}(z_{in}, z_{iUS})$ to 1), the smaller the bias of the least-squares slopes in (13); and (ii) the least-squares slopes in (13) are biased towards zero (attenuated) for all countries n , as long as the technological industry characteristics of all countries are positively correlated with those of the US ($\text{corr}(z_{in}, z_{iUS}) \geq 0$ for all n). Hence, as long as $\text{corr}(z_{in}, z_{iUS}) \geq 0$ for all countries n , the term in square brackets in (13) can be thought of as the so-called attenuation factor in the classical measurement error literature. This attenuation factor is larger—and hence the attenuation bias is smaller—for countries that are more similar technologically to the US.

That the country-specific least-squares slope estimates in (13) might be attenuated for all countries is not difficult to understand from the perspective of the classical measurement error literature, as US industry characteristics will generally proxy for industry characteristics of other countries with error. It is harder to see why, if all the slope estimates in (13) are attenuated, the standard benchmarking estimator may be subject to an amplification bias. This is possible because the attenuation bias of the least-squares slope estimates is heterogeneous across countries, with a smaller attenuation bias for countries that are more similar technologically to the US.

To see this point, it is useful to express the standard benchmarking estimator in (8) as a slope of slopes. We start from the least-squares slopes g_n in (13) obtained by regressing outcomes across industries on US industry characteristics separately for each country n . These country-specific slopes g_n are then regressed on the country characteristics x_n across countries. The least-squares slope of the second, cross-country regression is the standard benchmarking estimator in (8). To see this, note that

$$\begin{aligned} \frac{\sum_{n=1}^N g_n (x_n - \bar{x})}{\sum_{n=1}^N (x_n - \bar{x})^2} &= \beta \left(\frac{\sum_{n=1}^N \left[\left(1 - \frac{\sigma^2}{\sigma_{US}^2} \right) + \left(\frac{\sigma^2}{\sigma_{US}^2} \right) \rho_{nUS} \right] \gamma_n (x_n - \bar{x})}{\sum_{n=1}^N (x_n - \bar{x})^2} \right) \\ &= \beta \left[\left(1 - \frac{\sigma^2}{\sigma_{US}^2} \right) + \left(\frac{\sigma^2}{\sigma_{US}^2} \right) B \right] = b. \end{aligned} \quad (14)$$

The left-most expression in (14) is the standard expression for the slope of a least-squares regression, in this case of g_n on x_n . The first equality follows from substituting the least-squares slopes in (13) for g_n . The second equality uses (12) and the definition of B in (10), and the last equality uses the expression for b in (8) for the case $\alpha = 0$. The key message of the slope-of-slopes expression for the standard benchmarking estimator in (14) is that the bias of the estimator reflects how the attenuation factor of the country-specific least-squares slopes in (13) covaries with the country characteristics x_n . The amplification bias can emerge when the attenuation factor (bias) is larger (smaller) for countries with greater x_n .

We now illustrate the amplification bias in the simplest version of our framework.

The Amplification Bias in the Simplest Setting. The source of the amplification bias emerges most clearly when there are two different groups of countries and countries in the same group are identical. In this two-group setting, the formula for the benchmarking estimator in (14) simplifies to

$$b = \frac{g_S - g_D}{x_S - x_D} \quad (15)$$

where g_S and g_D are the country-specific slope estimates in (13) for countries in group S and group D , and x_S and x_D are the x -characteristics of countries in the two groups.

Now suppose that the US is part of group S . As countries in the same group are identical, this implies that all countries n in group S are identical technologically to the US, $\rho_{nUS} = 1$. As a result, (13) implies that the estimated country slopes and the true country slopes are the same for all countries in group S : $g_S = \gamma_S$. This is unsurprising as using US technological industry characteristics as a proxy for the technological industry characteristics of other countries in group S does not involve any measurement error.

On the other hand, suppose that countries in group D are technologically somewhat different from the US. The simplest approach is to think of these countries as having specific industry characteristics that are uncorrelated with US-specific industry characteristics. That is, $\rho_{nUS} = 0$ for all countries n in group D . If there is some cross-country heterogeneity in technological industry characteristics ($\sigma^2 > 0$), (13) implies that the estimated country slopes for all countries in group D are an attenuated version of the true slopes: $g_D = (1 - \sigma^2 / \sigma_{US}^2) \gamma_D$. This is because US industry characteristics are a noisy proxy for the technological industry characteristics of countries in group D .

Substituting the expressions for g_S and g_D we just obtained into (15) and using (12) yields

$$b = \beta \left[1 + \left(\frac{\sigma^2}{\sigma_{US}^2} \right) \frac{x_D}{x_S - x_D} \right]. \quad (16)$$

Hence, there will be an amplification bias, $|b| > |\beta|$ and $sign(b) = sign(\beta)$, if $x_S > x_D > 0$.

The bias can be large if the two groups of countries have very similar x -characteristics. This is because in this case, there is a strong positive association between the country characteristic x_n and technological similarity with the US.

Figure 1: The amplification bias in the simplest possible case.

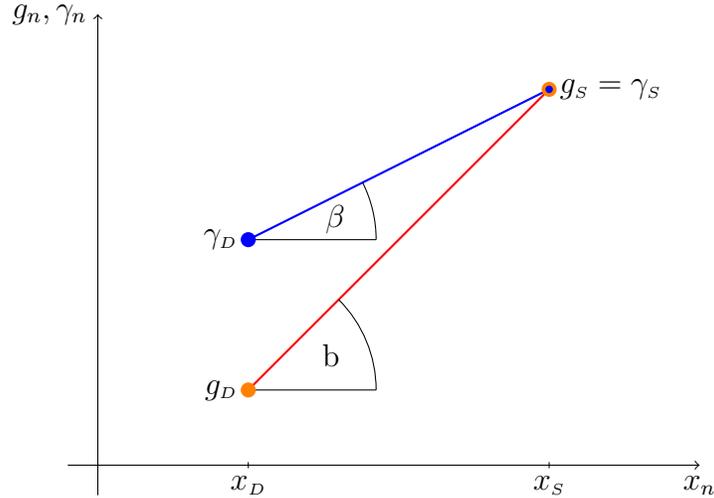


Figure 1 is a graphical illustration of the amplification bias in the two-group setting for $\beta > 0$. The two blue dots plot the true country-specific slopes γ_S and γ_D against x_S and x_D . The parameter β we want to estimate is the slope of the blue line connecting the two blue dots as (12) implies $\beta = (\gamma_S - \gamma_D)/(x_S - x_D)$. The two red dots plot the least-squares slope estimates g_S and g_D against x_S and x_D . Equation (15) implies that the benchmarking estimator b is the slope of the red line connecting the two red dots, $b = (g_S - g_D)/(x_S - x_D)$. The amplification bias $b > \beta > 0$ emerges because:

- (i) Countries in group S with high x -values have the same technological industry characteristics as the US, and US industry characteristics are therefore a perfect proxy for the industry characteristics of all high- x countries. Hence, there is no measurement error when the US is used to proxy for the industry characteristics of these countries. This implies that the least-squares slope estimates for these countries are equal to the true slopes, $g_S = \gamma_S$. That is, the blue and the red dot lie on top of each other.
- (ii) Countries in group D with low x -values have technological industry characteristics that are somewhat different from those of the US, and US industry characteristics therefore proxy for the technological industry characteristics of all low- x countries with some error. Hence, the least-squares slopes estimates g_D for these countries underestimates the true slopes, $g_D < \gamma_D$. That is, the red dot lies below the blue dot.

Hence, cross-country heterogeneity in technological industry characteristics implies that using the US industry proxy yields a consistent estimate of γ_S for high- x countries that are

technologically identical to the US, but a downwards biased estimate of γ_D for low- x countries that are technologically different from the US. Because the standard benchmarking estimator b is the slope of the red line connecting the red dots while the parameter of interest β is the slope of the blue line connecting the blue dots, this leads to an amplification bias, $0 < \beta < b$. More generally, the amplification bias of the standard benchmarking estimator arises when greater technological similarity between high- x countries and the US leads to a sufficiently smaller attenuation bias for the country-specific slope estimates of high- x countries.

3.2.3 The Bias of the Standard Benchmarking Estimator: the General Case

To characterize the bias of the standard benchmarking estimator more generally, it is useful to distinguish the case $\beta = 0$ and the case $\beta \neq 0$.

If $\beta = 0$, (8) simplifies to $b = \alpha A \sigma^2 / \sigma_{US}^2$ with $\sigma^2 / \sigma_{US}^2 < 1$. Hence, with cross-country heterogeneity in technological industry characteristics, $\sigma^2 > 0$, the standard benchmarking estimator is biased upwards if $\alpha A > 0$ and is biased downwards if $\alpha A < 0$.

If $\beta \neq 0$, the standard benchmarking estimator in (8) can be written as

$$b = \beta \left[\left(1 - \frac{\sigma^2}{\sigma_{US}^2} \right) + \left(\frac{\sigma^2}{\sigma_{US}^2} \right) \delta \right] \quad (17)$$

where $\sigma^2 / \sigma_{US}^2 < 1$ and δ is a function of A and B in (9)–(10)

$$\delta = \theta A + B \quad (18)$$

with

$$\theta = \frac{\alpha}{\beta}. \quad (19)$$

Hence, when there is cross-country heterogeneity in technological industry characteristics, $\sigma^2 > 0$, the bias of the standard benchmarking estimator depends on δ . If $\delta = 0$, the standard benchmarking estimator is attenuated. For example, our framework yields $\delta = 0$ when country-specific industry characteristics are uncorrelated across countries. If $\delta > 0$, there is a countervailing force that can weaken the attenuation bias or result in an amplification bias. If $\delta < 0$, the standard benchmarking estimates may have the wrong sign.

We now summarize how the bias of the standard benchmarking estimator depends on δ .

Proposition 1. [Bias of standard benchmarking estimator when $\beta \neq 0$]

1. If $0 \leq \delta \leq 1$, the standard benchmarking estimator is subject to an attenuation bias: b has the same sign as β but is biased towards zero, $sign(b) = sign(\beta)$ and $|b| \leq |\beta|$.
2. If $\delta > 1$, the standard benchmarking estimator is subject to an amplification bias: b has the same sign as β but is biased away from zero, $sign(b) = sign(\beta)$ and $|b| > |\beta|$.

3. If $\delta < 0$, the standard benchmarking estimator may be subject to an attenuation bias, an amplification bias, or may have a different sign than β , depending on σ^2/σ_{US}^2 .

4 Identification of β

We have seen that the standard benchmarking estimator used in the cross-industry cross-country literature does not identify the effect of interest when there is cross-country heterogeneity in technological industry characteristics ($\sigma^2 > 0$). Moreover, the bias cannot be signed if technologically similar countries are similar in terms of other characteristics ($A \neq 0$ or $B \neq 0$). We now examine how the effect of interest can be identified when there is cross-country heterogeneity in technological industry characteristics.

To get a first idea how the effect of interest might be identified and where the challenges lie, we return to the expression for the benchmarking estimator in (17). Inverting it yields $\beta = b/[1 + (\delta - 1)\sigma^2/\sigma_{US}^2]$. The right-hand-side parameter b can be identified using the standard benchmarking approach in the literature, and the variance of the US industry characteristics σ_{US}^2 is observable. If we can identify δ and σ^2 , we can therefore identify β . As we will show, δ can be identified from the variances and covariances of industry outcomes for different country pairs. If these variances and covariances would also identify the variance of country-specific industry characteristics σ^2 , identification of β would be straightforward. But the variances and covariances of industry outcomes do not identify σ^2 .

To see how the variances and covariances of industry outcomes for different country pairs help to identify β , we rewrite the model in (1) as

$$y_{in} = v_i + v_n + \gamma_i x_n + u_{in} \quad (20)$$

where

$$\gamma_i = \beta z_i \quad (21)$$

and

$$u_{in} = (\alpha + \beta x_n)\varepsilon_{in} \quad (22)$$

and v_i and v_n denote industry and country fixed effects.¹³ The industry-specific slopes γ_i capture the effect of the country characteristic on outcomes in different industries.

The effect of (unobservable) country-specific technological industry characteristics ε_{in} on industry outcomes is captured by u_{in} in (22). $E(u_{in}u_{im})$, the variances and covariances of u_{in} for industry i and countries n, m , reflect the effect of cross-country heterogeneity in

¹³These industry and country fixed effects capture the industry and country fixed effects in v_{in} and absorb αz_i in the industry fixed effect and z_n in the country fixed effect.

technological industry characteristics on the variances and covariances of industry outcomes in countries n, m . As a result, they play a central role for the identification of δ and β .

To see this, note that (3) and (22) imply that the variances and covariances $E(u_{in}u_{im})$ are

$$E(u_{in}u_{im}) = (\alpha\sigma + \beta\sigma x_n)(\alpha\sigma + \beta\sigma x_m)\rho_{nm} = \omega_{nm}. \quad (23)$$

That the ω_{nm} may allow us to identify δ is quite straightforward. From (18) and A and B in (9)–(10), it can be seen that δ depends on the ρ_{nm} , which capture how similar any two countries are technologically, and on α/β , which captures the direct effect of technological industry characteristics on industry outcomes relative to the industry-country-interaction effect. We should be able to infer these parameters entering δ from the ω_{nm} under some conditions as according to (23), the ω_{nm} depend on the ρ_{nm} ; on $\alpha\sigma$, which captures the direct effect of country-specific heterogeneity in technological industry characteristics on industry outcomes; and on $\beta\sigma$, which captures the industry-country-interaction effect of country-specific heterogeneity in technological industry characteristics on industry outcomes.

However, it can also be seen that the ω_{nm} in (23) will not allow us to identify the variance of country-specific industry characteristics σ^2 . This is because the ω_{nm} solely reflect σ^2 through its effects on outcomes, which is why σ only appears multiplied by either α or β . This is what makes the identification of β challenging.

To see when and how β can be identified, we now proceed in two steps. We first examine the identification of β for known ω_{nm} . Then we discuss how the ω_{nm} can be identified.

4.1 Identification of β for Known Ω

It is convenient to collect the variances and covariances ω_{nm} in (23) for all countries n, m in the $N \times N$ variance-covariance matrix Ω . The straightforward part of identification of β for known Ω is determining whether or not $\beta = 0$. The elements on the diagonal of Ω are equal to $\omega_{nn} = (\alpha\sigma + \beta\sigma x_n)^2$ for all countries n . As long as there is some cross-country heterogeneity in technological industry characteristics, $\sigma^2 > 0$, the ω_{nn} are independent of country characteristics if and only if $\beta = 0$. Hence, we obtain that $\beta = 0$ if the ω_{nn} are independent of x_n . On the other hand, $\beta \neq 0$ if the ω_{nn} depend on x_n .

The next question is how to identify β if the ω_{nm} depend on the country characteristics x_n . We first explain how Ω can be used to obtain two key parameters for the identification of β , namely δ and $(\beta\sigma)^2$. Then we show how δ and $(\beta\sigma)^2$ can be used to identify β .

Obtaining δ and $(\beta\sigma)^2$ from Ω is simple. We start by determining $\alpha\sigma$ and $\beta\sigma$ —and hence $(\beta\sigma)^2$ —from the variances $\omega_{nn} = (\alpha\sigma + \beta\sigma x_n)^2$. This is possible if there are at least two countries with different x -values, so that we have at least two equations in the two unknowns

$\alpha\sigma$ and $\beta\sigma$.¹⁴ Then we invert the expression for the covariances ω_{nm} for $n \neq m$ in (23) to get $\rho_{nm} = \omega_{nm}/[(\alpha\sigma + \beta\sigma x_n)(\alpha\sigma + \beta\sigma x_m)]$. This allows us to obtain the ρ_{nm} by combining the ω_{nm} with $\alpha\sigma$ and $\beta\sigma$. Once we have obtained $\alpha\sigma$, $\beta\sigma$, and the ρ_{nm} , it is straightforward to obtain A and B in (9)–(10), θ in (19), and hence $\delta = \theta A + B$ in (18).

To see when and how δ and $(\beta\sigma)^2$ obtained from $\mathbf{\Omega}$ allow us to identify β , we start from the expression for the bias of the standard benchmarking estimator $b - \beta = \beta(\delta - 1)\sigma^2/\sigma_{US}^2$ obtained by rearranging (17). Multiplying both sides of this equation by β yields $(b - \beta)\beta = (\delta - 1)(\beta\sigma)^2/\sigma_{US}^2$. The right-hand side parameters δ and $(\beta\sigma)^2$ can be obtained from $\mathbf{\Omega}$, and σ_{US}^2 is the observable variance of US industry characteristics. The parameter b is identified by the standard benchmarking approach in the literature. Hence, β is the only unknown of the quadratic equation

$$(b - \beta)\beta = \eta(\delta - 1) \quad (24)$$

where we defined

$$\eta = \frac{(\beta\sigma)^2}{\sigma_{US}^2}. \quad (25)$$

This establishes a key result: β is one of the solutions for q of the quadratic equation

$$(b - q)q = \eta(\delta - 1). \quad (26)$$

Generally, the quadratic equation in (26) has two solutions. In addition to the solution $q_1 = \beta$, there is a second solution $q_2 = \beta(\delta - 1)\sigma^2/\sigma_{US}^2$. We therefore need to analyze when we can determine which of the two solutions for q in (26) identifies β .

We start with the simplest case, which is when δ is positive and smaller than 2. In this case $\sigma^2/\sigma_{US}^2 < 1$ implies $(\delta - 1)\sigma^2/\sigma_{US}^2 \in (-1, 1)$. As the two solutions for q in (26) are $q_1 = \beta$ and $q_2 = \beta(\delta - 1)\sigma^2/\sigma_{US}^2$, this yields that β can be identified as the solution for q in (26) with the larger absolute value, $\beta = \max(|q_1|, |q_2|)$.

This is the simplest expression for β when δ is positive and smaller than 2. But the expression does not generalize to other cases where β is exactly identified. An alternative expression that holds for all cases where β is exactly identified is $\beta = \kappa b$, where b is the standard benchmarking estimator and κ is a function of the two solutions for q in (26)

$$\kappa = \max\left(\frac{q_1}{q_1 + q_2}, \frac{q_2}{q_1 + q_2}\right). \quad (27)$$

The next proposition, which is proven in the Appendix, summarizes this result.

Proposition 2. [Identifying β : sufficient condition in terms of identifiable δ] If

¹⁴There is no gain of using more than two ω_{nm} equations as additional equations leave results unchanged. When we use our identification results for estimation, we use all ω_{nm} equations of course.

$\delta \in [0, 2]$, β can be identified as $\beta = \kappa b$ where b is the probability limit of the standard benchmarking estimator and κ is defined in (27).

The next proposition gives a necessary and sufficient condition for the exact identification of β for known Ω .

Proposition 3. [Identifying β : necessary and sufficient condition in terms of identifiable δ and κ]

The effect of interest β can be exactly identified if and only if

$$\begin{aligned} & \text{either } \delta \geq 0 \quad \text{and} \quad \kappa \geq \frac{\delta-1}{\delta} \\ & \text{or} \quad \delta < 0 \quad \text{and} \quad \kappa \leq \frac{\delta-1}{\delta} \end{aligned} \tag{28}$$

where δ is defined in (18) and κ is defined in (27). If this condition is not satisfied, β is equal to one of the two solutions for q in (26), but it cannot be determined which.

When β is exactly identified, it can be obtained as

$$\beta = \kappa b \tag{29}$$

where b is the probability limit of the standard benchmarking estimator.

The proposition is proven in the Appendix. The idea is the following. The two solutions for q of the quadratic equation in (26) yield two candidate solutions for β . Each of these two candidate solutions can be combined with the variance of US industry characteristics and the identifiable parameter η in (25) to yield two candidate solutions for the country-specific technological heterogeneity parameter σ^2 . As at least some of the variation in technological industry characteristics reflects a global component, it must be that $0 \leq \sigma^2 < \sigma_{US}^2$. It turns out that this restriction is only satisfied by one of the two candidate solutions for σ^2 if the condition in (28) holds. Hence, only one of the two candidate solutions for β is consistent with the model and this solution is $\beta = \kappa b$. On the other hand, if the condition in (28) fails, both candidate solutions for β imply candidate solutions for σ^2 that are positive and smaller than σ_{US}^2 . As a result, both candidate solutions are consistent with the model and it is impossible to say which of the two solutions of (26) identifies β .

The necessary and sufficient condition in Proposition 3 is not easily interpreted in terms of the parameters of the underlying model. The next proposition gives the necessary and sufficient condition for identification in terms of σ^2 and δ .

Proposition 4. [Identifying β : necessary and sufficient condition in terms of model parameters] β can be exactly identified if and only if

$$(\delta - 1)^2 \left(\frac{\sigma^2}{\sigma_{US}^2} \right) \leq 1. \tag{30}$$

If this condition is not satisfied, β is one of the two solutions for q in (26), but it cannot be determined which.

Intuitively, Proposition 4 implies that β can be identified exactly if cross-country heterogeneity in technological industry characteristics is not too large (σ^2/σ_{US}^2 not too large) and/or if the association between countries' technological similarity with the US and their x -characteristics is not too strong (δ not too large in absolute value). On the other hand, if there is substantial cross-country heterogeneity in technological industry characteristics and/or countries' technological similarity with the US is strongly associated with their x -characteristics, it cannot be established which of the two solutions of (26) identifies β .

When exact identification of β is impossible, one could report both solutions for q in (26) as possible values for β . An alternative is to establish bounds on β in terms of the standard benchmarking estimator b . For $\delta > 2$, we have already established upper and lower bounds in Proposition 1. The next proposition establishes somewhat tighter bounds under the condition that $\delta > 2$ and that exact identification of β is impossible. For completeness, the proposition also gives bounds for the case $\delta < 0$ even though these are less useful. The proof of the proposition is in the Appendix.

Proposition 5. [Bounds on β] If the condition in (28) does not hold and exact identification of β is impossible, then

$$\begin{aligned} \text{if } \delta > 2 \text{ then } \frac{\beta}{b} &\in \left(\frac{1}{\delta}, \frac{\delta-1}{\delta}\right) \\ \text{if } \delta < 0 \text{ then } \frac{\beta}{b} &\notin \left[\frac{1}{\delta}, \frac{\delta-1}{\delta}\right]. \end{aligned} \tag{31}$$

For example, suppose that $\delta = 2.5$, b is positive, and (28) does not hold. In this case Proposition 5 implies that β is between $0.4b$ and $0.6b$. Hence, we can infer the range and the sign of the parameter of interest β from the standard benchmark estimator b . As another example, suppose that $\delta = -2.5$, b is positive, and (28) does not hold. Proposition 5 then implies that β is smaller than $-0.4b$ or larger than $0.6b$. Hence, we cannot establish an upper or lower bound for β , nor can we infer the sign of β from the sign of b .

4.2 Identification of Ω

Now that we have shown when and how β can be identified for known variance-covariance matrix Ω , we turn to the identification of Ω . Our approach is closely related to the identification of variance-covariance matrices in general least squares theory. In particular, the first step to identify Ω consists of least-squares estimation and the second step involves

understanding when and how the least-squares residuals can be used to identify Ω .¹⁵

The starting point to identify Ω is least-squares estimation of the model in (20). The least-squares residuals $\hat{u}_{in} = y_{in} - \hat{v}_i - \hat{v}_n - \hat{\gamma}_i x_n$, with hats denoting least-squares estimates, allow us to estimate $\frac{1}{I} \sum_{i=1}^I \hat{u}_{in} \hat{u}_{im}$ for all country pairs. These estimated variances and covariances depend on the ω_{nm} we collected in the variance-covariance matrix Ω and can therefore be used to identify Ω under some conditions.

Relating Ω to the variances and covariances of the residuals across industries.

We now derive the relationship between the variances and covariances across industries of the residuals \hat{u}_{in} for all pairs of countries, $\frac{1}{I} \sum_{i=1}^I \hat{u}_{in} \hat{u}_{im}$, and the elements ω_{nm} of Ω . The first step is to express the least-squares residuals \hat{u}_{in} in terms of the underlying disturbances u_{in} in (20)

$$\hat{u}_{in} = v_{in} - (x_n - \bar{x}) \sum_{k=1}^N \psi_k v_{ik} \quad (32)$$

where the v_{in} are the demeaned versions of u_{in}

$$v_{in} = u_{in} - \frac{1}{N} \sum_{m=1}^N u_{im} - \frac{1}{I} \sum_{j=1}^I u_{jn} + \frac{1}{N} \frac{1}{I} \sum_{m=1}^N \sum_{j=1}^I u_{jm} \quad (33)$$

and the ψ_k are the least-squares regression weights

$$\psi_k = \frac{x_k - \bar{x}}{\sum_{p=1}^N (x_p - \bar{x})^2}. \quad (34)$$

The second step is to calculate the probability limit as the number of industries goes to infinity of the variances and covariances of the residuals across industries for all country pairs, which we refer to as π_{nm}

$$\pi_{nm} = \text{plim}_{I \rightarrow \infty} \frac{1}{I} \sum_{i=1}^I \hat{u}_{in} \hat{u}_{im}. \quad (35)$$

We show in the Appendix that using (32)-(33) in (35) yields the following equations linking

¹⁵The main difference with GLS analysis is that we are interested in the variance-covariance matrix of the disturbances in (20) but not other model parameters, like the industry slopes for example. GLS analysis is generally interested in variance-covariance matrices because of their role in the efficient estimation of other model parameters. The reason we are not interested in the other model parameters in (20) is that they not help to identify β . For example, while the industry slopes depend on β , they also depend on the unobservable variation in the global component of technological industry characteristics.

π_{nm} and the elements ω_{nm} of $\mathbf{\Omega}$

$$\pi_{nm} = \omega_{nm} - \mu_n - \mu_m - (x_n - \bar{x})\lambda_m - (x_m - \bar{x})\lambda_n \quad (36)$$

where μ_n and λ_n are functions of the ω_{nm} detailed in the Appendix and

$$0 = \sum_{n=1}^N \lambda_n. \quad (37)$$

These equations are the basis for the identification of the variance-covariance matrix $\mathbf{\Omega}$ from the least-squares residuals of (20)-(22).

A structure for $\mathbf{\Omega}$. It is well understood that the identification of the variance-covariance matrix $\mathbf{\Omega}$ is impossible for an arbitrary matrix as (36) and (37) has more unknowns than linearly independent equations (e.g., Amemiya, 1985).¹⁶ For identification to be possible, the empirical framework must put some structure on $\mathbf{\Omega}$. The structures used in the literature depend on the application (e.g., Amemiya, 1985; Wooldridge, 2002; Conley, 2010).

We chose a structure for $\mathbf{\Omega}$ that has the implicit structure in the cross-industry cross-country literature as a special case but allows for substantial deviations from this baseline. The implicit structure for $\mathbf{\Omega}$ in the cross-industry cross-country literature is that differences between the technological industry characteristics of a country and global technological industry characteristics are completely idiosyncratic to each country. This implies that the technological industry characteristics of different countries are related through the global component only. Or put differently, the country-specific technological industry characteristics ε_{in} in (2) for any pair of countries $n \neq m$ are uncorrelated, i.e. $\rho_{nm} = 0$. As we have seen above, and as acknowledged in the cross-industry cross-country literature, the standard benchmarking estimator is attenuated in this case.

We chose a structure for $\mathbf{\Omega}$ that follows the cross-industry cross-country literature in that the technological industry characteristics of *some country pairs* are related through the global component only. I.e. for some country pairs $n \neq m$, $\rho_{nm} = 0$. But for all other country pairs, we allow for an entirely arbitrary correlation ρ_{nm} between the country-specific technological industry characteristics.

Specifically, our structure for $\mathbf{\Omega}$:

- (i) Allows for an arbitrary correlation ρ_{nm} between the country-specific technological industry characteristics of two countries if they are sufficiently similar. Two countries are taken to be sufficiently similar if the distance between their x -characteristics is below a threshold τ . When we set large values for the threshold τ , many country pairs satisfy

¹⁶We show this in the Appendix.

$|x_n - x_m| \leq \tau$, and our structure for $\mathbf{\Omega}$ therefore allows for arbitrary correlations ρ_{nm} between the country-specific technological industry characteristics of many country pairs. Formally, for these country pairs, technological similarity as measured by $corr(z_{in}, z_{im})$ is $[Var(z_i) + \sigma^2 \rho_{nm}] / (Var(z_i) + \sigma^2)$. Hence, the technological industry characteristics of these country pairs are *not* assumed to be related through the global component only and can be related in arbitrary ways to all country characteristics.

- (ii) When the distance between the x -characteristics of a country pair exceeds the threshold τ , their country-specific industry characteristics are taken to be uncorrelated, $\rho_{nm} = 0$.¹⁷ $\rho_{nm} = 0$ implies that the technological industry characteristics of these country pairs *are* related through the global technological component only, as implicitly assumed for all country pairs in the cross-industry cross-country literature. Formally, technological similarity as measured by $corr(z_{in}, z_{im})$ for country pairs with $\rho_{nm} = 0$ is $Var(z_i) / (Var(z_i) + \sigma^2)$. By increasing τ , we can reduce the number of country pairs with $\rho_{nm} = 0$ and therefore deviate substantially from the implicit assumption in the cross-industry cross-country literature that $\rho_{nm} = 0$ for all country pairs $n \neq m$.

We refer to this structure for the variance-covariance matrix $\mathbf{\Omega}$ as $\mathbf{\Omega}^\tau$ to capture that it depends on the threshold τ . What makes this structure for $\mathbf{\Omega}$ interesting in our context is that it corresponds to the implicit structure in the cross-industry cross-country literature for $\tau = 0$. We can move away from this baseline quite continuously and substantially by increasing τ . Moreover, the structure does not impose any functional form on how the technological similarity of country pairs with $|x_n - x_m| \leq \tau$ depends on country characteristics.

The size of the threshold τ must be interpreted relative to the distribution of the x -characteristic across countries. It is therefore often easier to think about the fraction of unrestricted ρ_{nm} implied by a threshold τ . For example, when τ is chosen very small, the fraction of unrestricted ρ_{nm} will be small as few country pairs will satisfy $|x_n - x_m| \leq \tau$. As a result, the assumed structure for $\mathbf{\Omega}$ will be similar to the implicit structure in the cross-industry cross-country literature. On the other hand, when the threshold τ is chosen large, the fraction of unrestricted ρ_{nm} will be large as many country pairs will satisfy $|x_n - x_m| \leq \tau$. As a result, the structure for $\mathbf{\Omega}$ can deviate quite substantially from the implicit structure in the cross-industry cross-country literature. (If the threshold τ is chosen so large that all country pairs can have different ρ_{nm} , we are not imposing any structure on the variance-covariance matrix $\mathbf{\Omega}$ and identification is impossible.)

As the choice is difficult in practice, we vary the threshold τ over the whole range that permits identification of $\mathbf{\Omega}$. Put differently, we allow the fraction of unrestricted ρ_{nm} to vary

¹⁷The approach can be thought of as a cross-country analogue of so-called K-dependence in time-series econometrics, which allows for any correlation between random variables at t and T if $|t - T| \leq \tau$ but assumes independence if $|t - T| > \tau$ (e.g., Amemiya, 1985).

between zero and the maximum that still permits identification of Ω . As this maximum can be surprisingly large, our structure for Ω can deviate substantially from the implicit structure in the cross-industry cross-country. In some cases, Ω can be identified for values of τ that leave 90% of the ρ_{nm} unrestricted. This amounts to little structure being put on the cross-country heterogeneity in technological industry characteristics. By varying the fraction of the unrestricted ρ_{nm} between zero and the maximum that permits identification, we can examine how sensitive the results for β are to the restrictions put on Ω .

Of course, other, more parsimonious structures for Ω could be chosen (and would generally be simpler to deal with). For example, the structures used in spatial econometrics for spatial dependence could be adapted to capture the technological similarity of countries as a function of their x -characteristics and other country characteristics (e.g., Conley, 2010).

Summarizing, we assume that if countries have sufficiently similar x -characteristics $|x_n - x_m| < \tau$, ρ_{nm} is unrestricted. On the other hand, $\rho_{nm} = 0$ if $|x_n - x_m| \geq \tau$. The threshold τ is set by us and we present results for the largest possible range allowing for the identification of Ω . Larger τ translate into a greater fraction of ρ_{nm} that are unrestricted.

A condition for identification of Ω . The structure Ω^τ for the variance-covariance matrix Ω assumes $\rho_{nm} = 0$ and hence $\omega_{nm} = 0$ in (23) for all country pairs with relatively different x -characteristics, $|x_n - x_m| \geq \tau$. We denote the number of such country pairs by Q . For these Q country pairs, (36) simplifies to

$$\pi_{nm} = -\mu_n^\tau - \mu_m^\tau - (x_n - \bar{x})\lambda_n^\tau - (x_m - \bar{x})\lambda_m^\tau. \quad (38)$$

These equations are the starting point for the identification of Ω^τ from the π_{nm} . In particular, we use these equations to try and determine the μ_n^τ and λ_n^τ for all n . Then we use (36) to determine the ω_{nm}^τ for all country pairs with relatively similar x -characteristics.

To take the first step and determine μ_n^τ and λ_n^τ , it is useful write the Q equations in (38) and the restriction in (37) in normal form

$$\boldsymbol{\pi} = \mathbf{G}^\tau \begin{pmatrix} \boldsymbol{\mu}^\tau \\ \boldsymbol{\lambda}^\tau \end{pmatrix} \quad (39)$$

where $\boldsymbol{\mu}^\tau = (\mu_1^\tau, \dots, \mu_N^\tau)'$ and $\boldsymbol{\lambda}^\tau = (\lambda_1^\tau, \dots, \lambda_N^\tau)'$ collect the $2N$ unknowns; $\boldsymbol{\pi}$ is a column vector of length $Q+1$ that collects the values on the left-hand side of equations (37) and (38); and \mathbf{G}^τ is a $(Q+1) \times 2N$ matrix of coefficients implied by the right-hand side of equations (37) and (38). By writing the equations in (37) and (38) in normal form, it becomes clear that $\boldsymbol{\mu}^\tau$ and $\boldsymbol{\lambda}^\tau$ can be determined if the matrix \mathbf{G}^τ has full rank.

An illustration of the identification condition. We can identify the variance-covariance matrix $\mathbf{\Omega}^\tau$ if the matrix \mathbf{G}^τ has full rank. This depends on the distance threshold τ and the distribution of the x -values across countries.

Table 1 illustrates this for three types of distributions for the x -values across countries. For each distribution, we draw x -values for 150 countries.¹⁸ We repeat this 300 times. For each draw we calculate the value for the maximum threshold τ such that \mathbf{G}^τ has full rank for all smaller τ . We refer to this value as τ_{\max} . As this value is somewhat difficult to interpret, we do two things to put it into perspective:

- (i) We calculate the average distance $|x_n - x_m|$ across all possible country pairs for each draw. This allows comparing τ_{\max} with the average distance in the x -characteristics across all country pairs and get a sense whether τ_{\max} is relatively large or small.
- (ii) We calculate the number of countries with unrestricted ρ_{nm} that are implied by τ_{\max} . We then report this number relative to the total number of country pairs. For example, if this ratio is 0.8, the ρ_{nm} are unrestricted for 80% of all country pairs.

Table 1 reports these statistics averaged across the 300 draws we take. We first present results for the case where the country characteristics are uniformly distributed between 0 and 1. The distance $|x_n - x_m|$ averaged across all country pairs is 0.33. The maximum value of the distance threshold τ that still permits identification (τ_{\max}) is 0.49. The number of country pairs with unrestricted ρ_{nm} relative to the total number of country pairs at τ_{\max} is 74%. The statistics in the last two columns remain nearly unchanged when we vary the support of the uniform distribution (not in the table).

As a second illustration, Table 1 shows results for the case where the country characteristics are drawn from a normal distribution with mean 0 and a standard deviation of 1. The distance $|x_n - x_m|$ averaged across all country pairs is 1.13. τ_{\max} is 2.42. The number of country pairs with unrestricted ρ_{nm} relative to the total number of country pairs at τ_{\max} is 93%. The statistics in the last two columns do not vary with the mean of the normal distribution and remain nearly unchanged when we vary the standard deviation (not in the table). The third illustration in Table 1 is for the exponential distribution and yields results similar to the normal distribution.

5 An Application

We now apply our identification results. We start by explaining how to go from identification to estimation. Then we use the approach to reestimate Nunn (2007).

¹⁸This is approximately the number of countries in our application of the new benchmarking estimator below. We obtain very similar results for 75 and 250 countries.

Table 1: Identification of the variance-covariance matrix.

Distribution	Average distance between x_n across all country pairs	Maximum threshold τ allowing identification (τ_{\max})	Country pairs with unrestricted ρ_{nm} relative to total number of country pairs at τ_{\max}
Uniform on $[0, 1]$	0.33	0.49	0.74
Standard normal	1.13	2.58	0.93
Exp. with $\lambda = 1$	1.01	2.55	0.91

5.1 From Identification to Estimation

We first explain how our identification results can be used to obtain consistent estimates of q in (26) in five steps. Once we have estimated the solutions for q , we estimate β using Proposition 2 or Proposition 3, or obtain bounds for β using Proposition 5.

Step 1: Estimate (20) with least squares and then use the residuals to estimate the variances and covariances across industries of the residuals for all country pairs

$$\hat{\pi}_{nm} = \frac{1}{I} \sum_{i=1}^I \hat{u}_{in} \hat{u}_{im}. \quad (40)$$

These variances and covariances are consistent estimators of the π_{nm} in (35) as the number of industries I goes to infinity.

Step 2: Estimate $\boldsymbol{\mu}^\tau$ and $\boldsymbol{\lambda}^\tau$ on the basis of (39). We start by obtaining the matrix \mathbf{G}^τ for different distance cutoffs τ . We begin with very small values of τ . If all countries have different x -characteristics (as in our application below), this implies that the $\rho_{nm} = 0$ condition is imposed for all country pairs $n \neq m$ and that $\boldsymbol{\Omega}$ is a diagonal matrix (as implicitly assumed in the cross-industry cross-country literature). The implied matrix \mathbf{G}^τ is of full rank. We then increase τ up to the maximum value still yielding a matrix \mathbf{G}^τ of full rank. To estimate $\boldsymbol{\mu}^\tau$ and $\boldsymbol{\lambda}^\tau$ on the basis of (39), we also need an estimator of the column vector $\boldsymbol{\pi}$. We obtain this estimator by replacing the π_{nm} collected in the vector $\boldsymbol{\pi}$ with the estimates $\hat{\pi}_{nm}$ in (40). Of course, we cannot estimate $\boldsymbol{\mu}^\tau$ and $\boldsymbol{\lambda}^\tau$ by simply replacing $\boldsymbol{\pi}$ with $\hat{\boldsymbol{\pi}}$ in (39). This is because generally $\hat{\boldsymbol{\pi}} \neq \boldsymbol{\pi}$ due to sampling error and the equation system in (39) would therefore be overdetermined. Instead, $\boldsymbol{\mu}^\tau$ and $\boldsymbol{\lambda}^\tau$ are estimated by applying

least squares to

$$\widehat{\boldsymbol{\pi}} = \mathbf{G}^\tau \begin{pmatrix} \boldsymbol{\mu}^\tau \\ \boldsymbol{\lambda}^\tau \end{pmatrix} + \mathbf{v} \quad (41)$$

where \mathbf{v} is a column vector of length $Q+1$ that captures the sampling error $\widehat{\boldsymbol{\pi}} - \boldsymbol{\pi}$. Because $\widehat{\boldsymbol{\pi}}$ is a consistent estimator of $\boldsymbol{\pi}$ as the number of industries I goes to infinity, the least-squares estimators $\widehat{\boldsymbol{\mu}}^\tau$ and $\widehat{\boldsymbol{\lambda}}^\tau$ are consistent estimators of $\boldsymbol{\mu}^\tau$ and $\boldsymbol{\lambda}^\tau$.

Step 3: Estimate the non-zero elements ω_{nm}^τ of $\boldsymbol{\Omega}^\tau$ by combining (36) with $\widehat{\boldsymbol{\mu}}^\tau$, $\widehat{\boldsymbol{\lambda}}^\tau$, and $\widehat{\boldsymbol{\pi}}$. This yields

$$\widehat{\omega}_{nm}^\tau = \widehat{\mu}_n^\tau + \widehat{\mu}_m^\tau + (x_n - \bar{x})\widehat{\lambda}_m^\tau + (x_m - \bar{x})\widehat{\lambda}_n^\tau + \widehat{\pi}_{nm}. \quad (42)$$

Consistency of the $\widehat{\omega}_{nm}^\tau$ follows from the consistency of $\widehat{\boldsymbol{\mu}}^\tau$, $\widehat{\boldsymbol{\lambda}}^\tau$, and $\widehat{\boldsymbol{\pi}}$. The estimates of ω_{nm}^τ allow us to estimate θ , $\beta\sigma$, and the nonzero ρ_{nm} . The estimates of θ and $\beta\sigma$ are obtained by combining the expressions for the variances $\omega_{nn} = (\theta + x_n)^2(\beta\sigma)^2$ in (23) with our estimates $\widehat{\omega}_{nn}^\tau$. This yields

$$\widehat{\omega}_{nn}^\tau = (\theta + x_n)^2(\beta\sigma)^2 + v_{nn} \quad (43)$$

where v_{nn} captures sampling error. The nonlinear least-squares estimates of θ and $\beta\sigma$ are then combined with our estimates of the nonzero ω_{nm}^τ for $n \neq m$ and the expression for the covariances in (23) to estimate the nonzero ρ_{nm}^τ using that $\rho_{nm}^\tau = \omega_{nm}^\tau / [(\theta + x_n)(\theta + x_m)(\beta\sigma)^2]$. Moreover, our estimate of $\beta\sigma$ can be combined with the variance of the industry characteristics in the benchmark country σ_{US}^2 to estimate $\widehat{\eta}$ using (25). Consistency follows from the consistency of the $\widehat{\omega}_{nm}^\tau$.

Step 4: Use the estimates $\widehat{\rho}_{nm}^\tau$ to estimate \widehat{A}^τ using (9) and \widehat{B}^τ using (10). Hence, we have all the elements to estimate $\widehat{\delta}^\tau$ using (18)

$$\widehat{\delta}^\tau = \widehat{\theta}\widehat{A}^\tau + \widehat{B}^\tau. \quad (44)$$

Step 5: Replace δ and η in (26) by the consistent estimates $\widehat{\delta}^\tau$ and $\widehat{\eta}$. This allows us to obtain consistent estimates of q by solving

$$(\widehat{b} - q)q = \widehat{\eta}^\tau(\widehat{\delta}^\tau - 1) \quad (45)$$

where \widehat{b} is the standard benchmarking estimator.

The estimates of q based on (45) can be used to estimate β as explained in Proposition 2 and Proposition 3 or to obtain bounds on β as explained in Proposition 5. Confidence bands

of all our estimates are obtained by bootstrapping.¹⁹

5.2 Reestimating Nunn (2007)

Nunn employs data on exports for up to 222 industries in up to 146 countries to show that institutional quality has a positive effect on comparative advantage in industries that depend more on relationship-specific intermediate inputs.²⁰ In terms of the model in (1), the institutional quality of countries takes the place of x_n and their log exports in industry i takes the place of y_{in} . The theoretically relevant industry characteristic z_{in} is the relationship-specific intermediate-input intensity of production. The benchmark country used to obtain proxies for how intensively industries use relationship-specific inputs is the US.

We apply the approach in the previous section to reestimate Nunn’s baseline specification without controls and his specification with controls for human and physical capital. We report our estimates of δ and β as a function of the threshold distance τ and the share of unrestricted, and hence estimated, correlation coefficients ρ_{nm} .

5.2.1 Results for the Baseline Specification

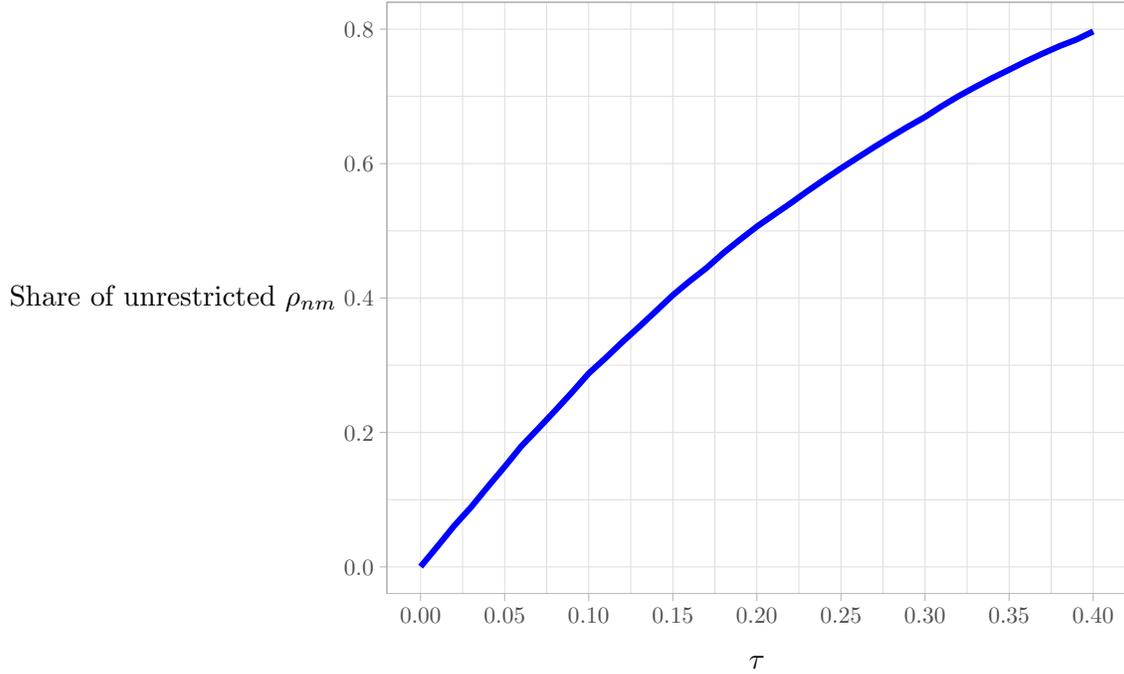
Figure 2 plots the threshold distance τ on the horizontal axis and the number of unrestricted correlation coefficients ρ_{nm} relative to the total number of country pairs on the vertical axis (there are 10,585 country pairs as Nunn has the necessary data for 146 countries). When τ is very small, the condition $\rho_{nm} = 0$ is assumed for all country pairs $n \neq m$ (all countries have different institutional quality in the Nunn data). Hence, the number of unrestricted ρ_{nm} relative to the total number of country pairs is 0. This corresponds to the implicit assumption in the cross-industry cross-country literature. For $\tau = 0.2$, around half the ρ_{nm} are unrestricted and must therefore be estimated. As τ goes to 0.4, 80% of the ρ_{nm} are unrestricted and must be estimated. For values of τ strictly larger than 0.4, \mathbf{G}^τ no longer has full rank and $\boldsymbol{\mu}^\tau$ and $\boldsymbol{\lambda}^\tau$ cannot be determined. Hence, $\boldsymbol{\Omega}^\tau$ cannot be identified.

Figure 3 summarizes our results for δ . Estimates are shown as green dots and 95% confidence intervals are marked as green lines. The area shaded in light grey marks values of δ that according to Proposition 1 result in an attenuation bias of the standard benchmarking estimator. The area shaded in darker grey marks values of δ that according to Proposition 1 result in an amplification bias of the standard benchmarking estimator. For very small values of τ , we obtain $\delta = 0$. This is unsurprising as the condition $\rho_{nm} = 0$ for $n \neq m$ is assumed

¹⁹Bootstrapping the confidence intervals of our estimates of δ and β involves reshuffling the u_{in} in (20) across industries for each country 300 times and each time reestimating $\boldsymbol{\pi}$, $\boldsymbol{\mu}$, $\boldsymbol{\lambda}$, ω_{nm} , θ , $\beta\sigma$, A , B , ρ_{nm} , δ , η , q_1 , q_2 , and β . The confidence intervals for our point estimates of δ and β are obtained from the distributions of the estimated δ and β .

²⁰See Levchenko (2007) and Costinot (2009) for related empirical and theoretical findings on the effect of institutional quality on comparative advantage.

Figure 2: Share of unrestricted ρ_{nm} as a function of τ for Nunn’s baseline specification.

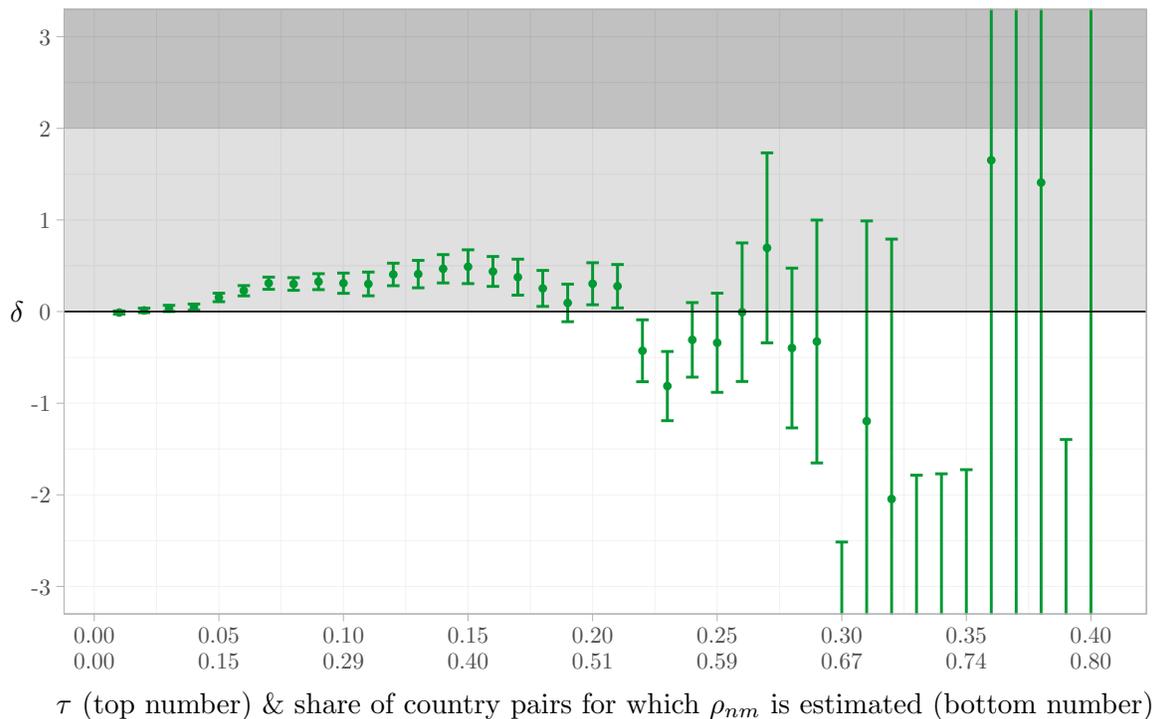


for all country pairs in this case (as implicitly assumed in the cross-industry cross-country literature). Hence, $A = B = 0$ in (9)–(10) and $\delta = 0$ in (18). Estimates of δ remain very small for values of τ smaller than 0.02. Point estimates are between -0.01 and $+0.01$ and 95% confidence intervals include 0. Hence, we cannot reject $\delta = 0$. According to Proposition 1, $\delta = 0$ implies that the standard benchmarking estimator is subject to an attenuation bias. According to Proposition 2, $\delta = 0$ implies that we can estimate β as $\beta = \kappa b$ with κ given in (27). Figure 4 shows our point estimates for β as orange dots and 95% confidence intervals as orange lines. Point estimates are around 7.2, around 10% larger than Nunn’s estimate of 6.6 obtained with the standard benchmarking estimator (marked by the horizontal black line).²¹ The 95% confidence intervals of our estimates are between 6.7 and 7.7.

For values of τ between 0.03 and 0.21, point estimates of δ in Figure 3 are between 0.04 and 0.49. The 95% confidence bands are strictly between 0 and 1, except for $\tau = 0.19$. The data therefore support values of δ greater than 0 but below 1. Proposition 1 implies that for $0 \leq \delta \leq 1$, the standard benchmarking estimator is subject to an attenuation bias. Proposition 2 implies that for $0 \leq \delta \leq 1$, we can estimate β as $\beta = \kappa b$. Figure 4 shows our estimates for β . Point estimates are between 7.2 and 6.8. Hence, the difference with Nunn’s benchmarking estimate of 6.6 is smaller than what we obtained for very small τ . This is because very small values for τ imply $\delta = 0$, and the bias of the standard benchmarking

²¹The estimate reported by Nunn differs because it is standardized. We report non-standardized estimates throughout.

Figure 3: Estimates of δ for Nunn’s baseline specification.



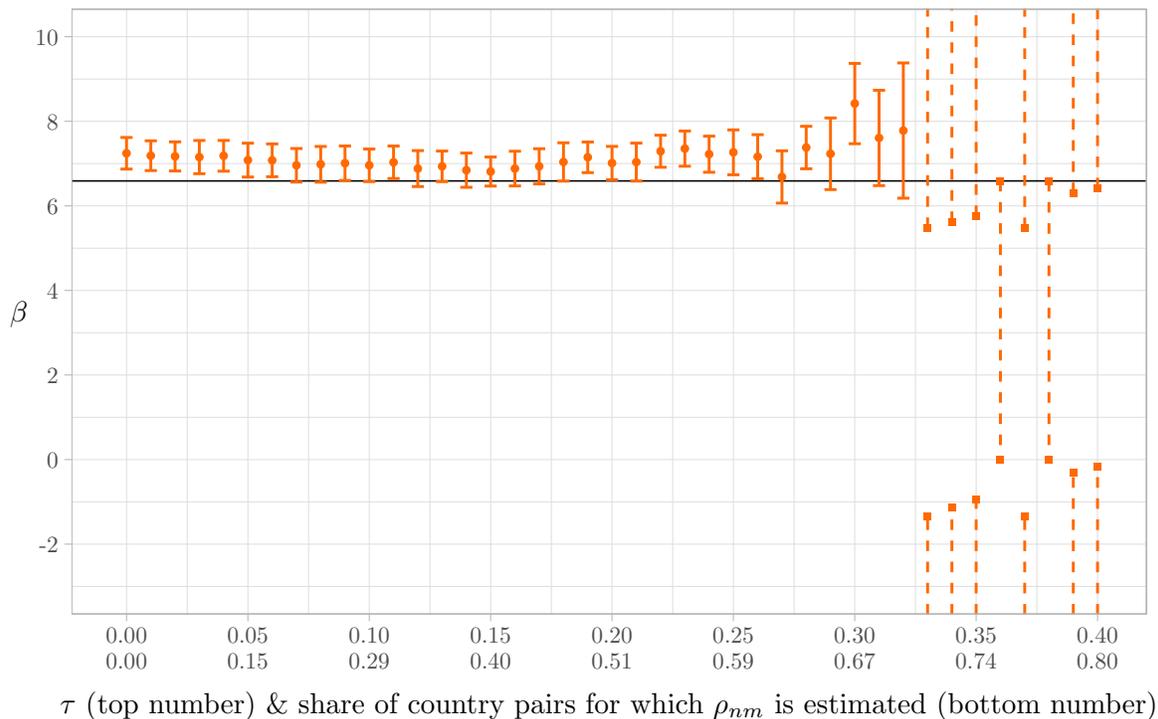
estimator is solely shaped by a force generating attenuation in this case. When $\delta > 0$, the bias of the standard benchmarking estimator is also shaped by a force countervailing the attenuation bias. The 95% confidence intervals of our estimates lie between 6.5 and 7.6.

For values of τ between 0.22 and 0.32, estimates of δ in Figure 3 are negative, except for $\tau = 0.27$. As a result, we cannot use Proposition 2 to estimate β . However, we can still estimate β as $\beta = \kappa b$ as our point estimates of κ satisfy the condition for exact identification in Proposition 3.²² Figure 4 shows our estimates for β . Point estimates are between 6.6 and 8.5 and therefore up to 30% larger than Nunn’s standard benchmarking estimate of 6.6 (the horizontal black line). The 95% confidence intervals of our estimates lie between 6 and 9.5.

For values of τ between 0.33 and 0.4, our estimates of δ in Figure 3 become very noisy. The range of 95% confidence intervals varies between 6 and 82 (we do not show the full intervals as this would make the figure unreadable). This likely reflects that for $\tau \geq 0.33$, the correlation coefficients ρ_{nm} of at least 70% of the 10,585 country pairs are being estimated. Point estimates of δ for τ between 0.33 and 0.4 are mostly negative. As the necessary and sufficient condition for exact identification of β in Proposition 3 is not satisfied, we can only establish the bounds in Proposition 5. Figure 4 illustrates the values of β consistent with

²²For simplicity, we are evaluating the condition in (28) based on the point estimates of δ and κ . A more complete approach would be to test the condition. To do so, note that the two cases in (28) can be combined in a single condition $\delta(\kappa - 1) + 1 \geq 0$. Bootstrapping the 95% confidence levels of the left-hand side of the inequality would allow testing this condition.

Figure 4: Estimates of β for Nunn’s baseline specification.



these bounds as dashed lines delimited by squares.

Overall, our estimation approach applied to Nunn’s baseline specification yields estimates that are close to Nunn’s even when we impose little structure on the cross-country heterogeneity in technological industry characteristics, i.e. as many as 70% of the ρ_{nm} are unrestricted and hence estimated. Sometimes this is because the countervailing forces generating an attenuation and amplification bias of the standard benchmarking estimator partly offset each other. Our estimates become very noisy and/or exact identification becomes impossible when more than 70% of the ρ_{nm} are unrestricted (the theoretical limit to identification is when 80% of the ρ_{nm} are unrestricted).

5.2.2 Results with Controls for Human and Physical Capital

Building on Romalis (2004) and other studies in international trade, Nunn also presents results controlling for the effect of human and physical capital on comparative advantage. He does so by augmenting his baseline specification with an interaction between country-level human capital and the human-capital-intensity of industries as well as an interaction between country-level physical capital and the physical-capital-intensity of industries.

We reestimate Nunn’s specification controlling for the effect of human and physical capital on comparative advantage using the new benchmarking estimator. The implementation follows the same steps as for Nunn’s baseline specification, except that least-squares estima-

tion of (20) accounts for the effect of human and physical capital following Nunn.

Figure 5: Share of unrestricted ρ_{nm} as a function of τ for Nunn’s specification with controls for human and physical capital.

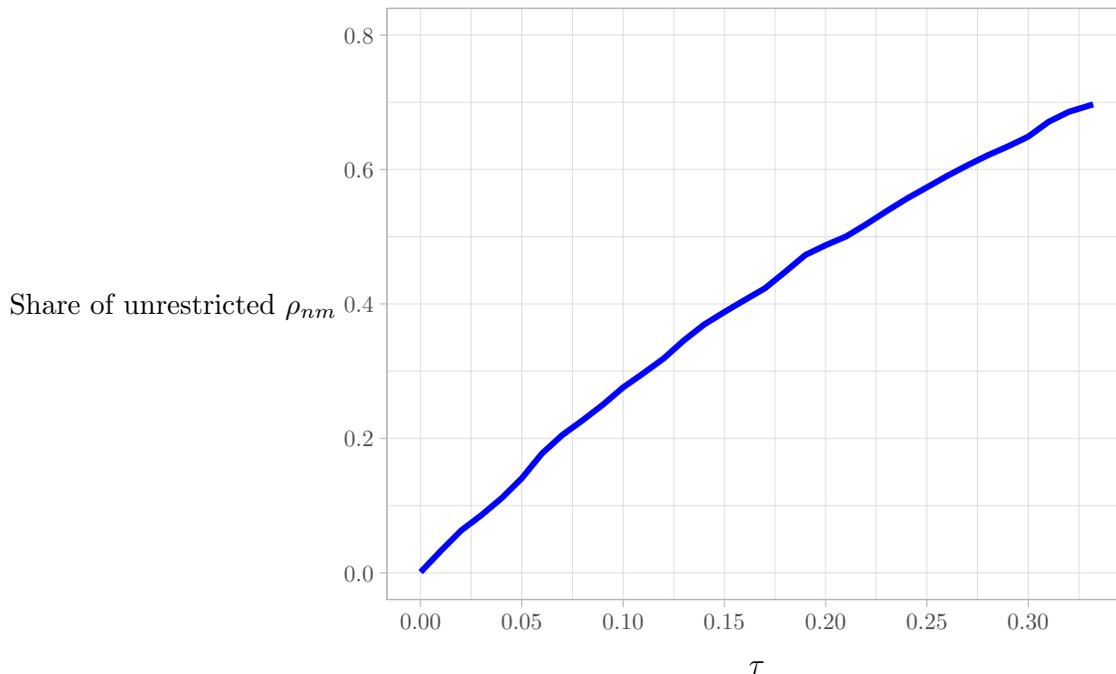
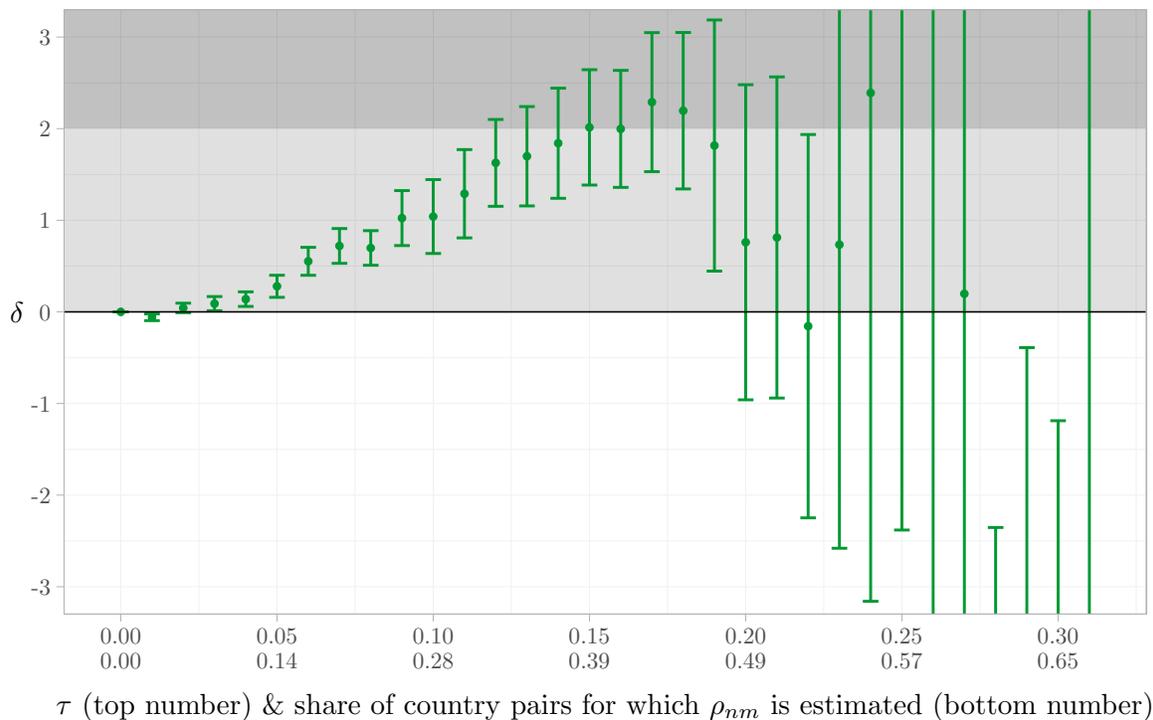


Figure 5 plots the threshold distance τ on the horizontal axis and the number of unrestricted correlation coefficients ρ_{nm} relative to the total number of country pairs on the vertical axis (there are only 2,415 country pairs in this specification, as Nunn has the necessary data for fewer countries). The ratio starts at 0 when τ is very small. This corresponds to the implicit assumption in the cross-industry cross-country literature. For $\tau = 0.2$, around half the ρ_{nm} are unrestricted and must therefore be estimated. As τ goes to 0.31, 70% the ρ_{nm} are unrestricted and must therefore be estimated. For larger values of τ , \mathbf{G}^τ no longer has full rank. Hence, $\mathbf{\Omega}^\tau$ can no longer be identified.

Figure 6 summarizes our results for δ . The area shaded in light grey continues to mark values of δ that according to Proposition 1 result in an attenuation bias of the standard benchmarking estimator. The area shaded in darker grey marks values of δ that according to Proposition 1 result in an amplification bias of the standard benchmarking estimator. For values of τ smaller than 0.02, point estimates of δ are between -0.01 and $+0.01$ and 95% confidence intervals include 0. Hence, we cannot reject $\delta = 0$. According to Proposition 1, $\delta = 0$ implies that the standard benchmarking estimator is attenuated. According to Proposition 2, $\delta = 0$ implies that we can estimate β as $\beta = \kappa b$. This yields estimates of β around 7, see Figure 7. These estimates are about 10% larger than Nunn’s point estimate of 6.4 obtained with the standard benchmarking estimator (marked by the horizontal black

Figure 6: Estimates of δ for Nunn’s specification with controls for human and physical capital.



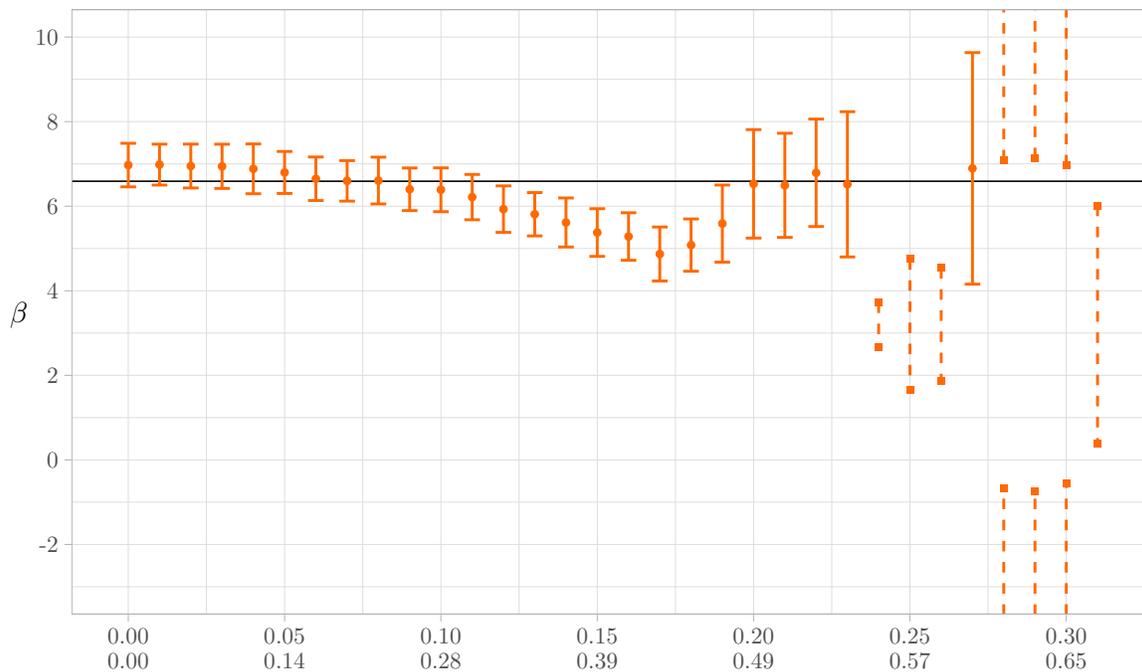
line). The 95% confidence intervals of our estimates for β lie between 6.4 and 7.5.

For values of the threshold distance τ between 0.03 and 0.11, point estimates of δ in Figure 6 are between 0.09 and 1.3. The 95% confidence intervals are between 0 and 2. Hence, the data support values of δ between 0 and 2. According to Proposition 2, $0 \leq \delta \leq 2$ implies that we can estimate β as $\beta = \kappa b$. This yields the estimates of β in Figure 7. These are sometimes above Nunn’s estimate of 6.4 and sometimes below. This makes sense as according to Proposition 1, the standard benchmarking estimator is subject to an attenuation bias when δ is between 0 and 1 and subject to an amplification bias when δ is greater than 1. The 95% confidence intervals of our estimates of β are between 5.7 and 7.5.

When τ is between 0.12 and 0.19, estimates of δ in Figure 6 are between 1.6 and 2.3. Hence, the standard benchmarking estimator is subject to an amplification bias according to Proposition 1. The condition for exact identification in Proposition 3 is always satisfied and we can therefore estimate β as $\beta = \kappa b$. Estimates of β in Figure 7 are between 5.9 and 4.9, up to 25% smaller than Nunn’s estimate of 6.4. The 95% confidence intervals of our estimates lie between 6.5 and 4.2.

For values of τ between 0.2 and 0.23, estimates of δ in Figure 6 are generally between 0 and 1. According to Proposition 2, $0 \leq \delta \leq 1$ implies that we can estimate β as $\beta = \kappa b$. Our point estimates of β in Figure 7 are between 6.5 and 6.7, only slightly larger than Nunn’s

Figure 7: Estimates of β for Nunn’s specification with controls for human and physical capital.



τ (top number) & share of country pairs for which ρ_{nm} is estimated (bottom number)

estimate of 6.4. The 95% confidence intervals of our estimates lie between 4.8 and 7.8.

For values of τ between 0.24 and 0.31, our estimates of δ are very noisy. The range of 95% confidence intervals varies between 11 and 42 (we do not show the full intervals as the figure would become unreadable). This likely reflects that we are approaching the limits of identification as for $\tau \geq 0.24$, the correlation coefficients ρ_{nm} of at least 55% of the 2,415 country pairs are being estimated.

Overall, our estimation approach applied to Nunn’s specification with controls for human and physical capital yields results that are similar to Nunn’s even when as many as 55% of the ρ_{nm} are unrestricted and hence estimated. When there are larger discrepancies between our estimates and those of Nunn, the forces generating an amplification bias of the standard benchmarking estimator dominate those generating an attenuation bias. As a result, our estimates tend to indicate smaller effects than Nunn’s. Our estimates become very noisy and/or exact identification becomes impossible when more than 55% of the ρ_{nm} are unrestricted (the theoretical limit to identification is when 70% of the ρ_{nm} are left unrestricted).

6 Conclusion

Using industry data to examine the economic effects of cross-country differences in financial development, institutional quality, human capital, and other potential determinants of aggregate economic activity is attractive for two main reasons. It permits testing whether the impact of, say, financial underdevelopment or malfunctioning institutions is strongest in the industries where it should be theoretically. This helps bringing empirical work closer to the mechanisms emphasized by economic theory. Moreover, using international industry data also allows addressing some of the reverse causation and omitted variable issues present in cross-country empirical work.

But the cross-industry cross-country approach is not without pitfalls. It requires specifying which industries should be affected most by financial underdevelopment, malfunctioning institutions, etc. As the relevant technological industry characteristics are unobservable in most countries, they are proxied by industry characteristics in a benchmark country. That this can lead to an attenuation bias is unsurprising and acknowledged in the literature.

What appears not understood in the cross-industry cross-country literature is that the estimation approach used can lead to an amplification bias or spurious results. Amplified or spurious estimates can arise when technologically similar countries are similar in other dimensions, like financial development, institutional quality, or human capital for example.

As the estimation approach used in the cross-industry cross-country literature does not identify the effect of interest in the literature when there is cross-country heterogeneity in technological industry characteristics, it is important to develop alternatives. To do so, we first provided an analysis of identification of the effect of interest. We then used the new identification approach to reestimate the effect of institutional quality on comparative advantage in industries that rely on relationship-specific inputs in Nunn (2007). Overall, our estimates tend to be similar to Nunn's even when we impose little structure on cross-country heterogeneity in technological industry characteristics.

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