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Prices, Plant Size, and Product Quality

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Drawing on uncommonly rich and representative data from the Colombian manufacturing census, this paper documents new empirical relationships between input prices, output prices, and plant size and proposes a model of endogenous input and output quality choices by heterogeneous firms to explain the observed patterns. The key empirical facts are that, on average within narrowly defined sectors, (1) larger plants charge more for their outputs and (2) larger plants pay more for their material inputs. The latter fact generalizes the well-known positive correlation between plant size and wages. Similar correlations hold between prices and export status. We show that the empirical patterns are consistent with a parsimonious extension of the Melitz (2003, "The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity," *Econometrica*, **71**, 1695–1725) framework to include endogenous choice of input and output quality. Using a measure of the scope for quality differentiation from Sutton (1998, *Technology and Market Structure: Theory and History*. Cambridge: MIT Press), we show that differences across sectors in the relationships between prices and plant size are consistent with our model. Available evidence suggests that differences in observable measures of market power do not provide a complete explanation for the empirical patterns. We interpret the results as supportive of the hypothesis that quality differences of both inputs and outputs play an important role in generating the price–plant size correlations.

Key words: Product prices, Quality, Plant size, Heterogeneous firms

JEL Codes: L11, F12, O14

1. INTRODUCTION

The increasing availability of micro-data on manufacturing plants has fueled a growing literature on heterogeneity at the plant level. Among the well-established stylized facts are that exporting plants are larger than non-exporters, that they have higher measured total factor productivity (TFP), and that they pay higher wages, even within narrowly defined sectors.[1] The Melitz (2003) model of heterogeneous firms under monopolistic competition has proven to be extremely useful in making sense of the new facts and has become the workhorse framework for analysing the behaviour of individual plants in an international context.[2]

[1] See Tybout (2003) for a review of work in this area.

[2] Bernard *et al.* (2003) provide an important alternative framework that is also consistent with many of the plant-level facts.

But while the first generation of plant-level data sets greatly expanded the research frontier, they typically contained little information on products, either goods produced ("outputs") or inputs purchased, and even less information on output or input prices. As a consequence, the heterogeneous-firm literature has had relatively little to say about heterogeneity at the product level. Empirical analyses (with a few notable exceptions, discussed below) have tended to be limited to plant-level variables such as revenues, average wages, and TFP. Theoretical models have tended to treat inputs and outputs in a highly stylized way. In particular, the Melitz model assumes that the lone input, labour, is homogeneous. In addition, although the model can be interpreted in terms of quality-differentiated outputs (as we discuss below), the interpretation of the model that has become standard treats outputs as symmetric, differentiated solely on a horizontal dimension.[3]

Drawing on a rich and nationally representative data set—the Colombian manufacturing census, which contains prices and physical quantities of all inputs and outputs used by all manufacturing plants with 10 or more workers over the period 1982–2005—this paper documents new empirical relationships between input prices, output prices, and plant size and shows that they are consistent with a parsimonious model of endogenous input and output quality choices by heterogeneous firms.[4] Observed differences across sectors in the relationships between prices and plant size are consistent with our model. Available evidence suggests that observable differences in market power across plants do not provide a complete explanation of the empirical patterns. We interpret the results as supportive of the hypothesis that quality differences of both inputs and outputs play an important role in generating the price–plant size correlations.

The paper proceeds in several steps. After describing the data and providing illustrative examples in the next section, we turn in Section 3 to documenting two stylized facts about plants and product prices. First, on average, there is a robust positive correlation between plant size (measured by sales or employment) and output prices: within narrow product categories, larger plants charge higher prices for the goods they produce. Second, and perhaps more surprisingly, there is a robust positive correlation between plant size and *input* prices: within narrow categories, larger plants pay more for the materials they purchase. This second fact generalizes the well-known positive correlation between plant size and wages—what Brown and Medoff (1989) refer to as the "employer size-wage effect." The novelty here is to show that this pattern applies not only to labour—the one input for which a price is observed in standard data sets—but on average to all material inputs. The correlations between prices and export status are similar to those between prices and plant size.

With these facts documented, in Section 4 we show that they are consistent with a parsimonious extension of the Melitz (2003) framework to include endogenous choice of input and output quality. As in Melitz (2003), *ex ante* homogeneous potential entrepreneurs pay a fixed cost to receive a "capability" draw[5] and are heterogeneous *ex post*. To the standard framework, we add a perfectly competitive but quality-differentiated intermediate-input sector. We consider two variants of the production function for output quality. In the first, input quality and plant capability are complements in generating output quality, and upgrading quality does not require fixed costs. In the second variant, closer in spirit to the seminal work of Sutton (1991, 1998, 2007), producing high-quality outputs requires both fixed costs and high-quality inputs, but there is no

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[3] A number of recent papers have considered multiproduct firms but have not focused on quality choices by firms (see *e.g.* Bernard, Redding and Schott (2010, forthcoming) and Eckel and Neary (2010)).

[4] In the Colombian data, we do not know which plants belong to which firms, and must conduct the analysis at the level of plants. Hereafter, we will treat plants as single-establishment firms, and use the terms plant and firm interchangeably.

[5] Melitz (2003) refers to this parameter as "productivity"; to avoid confusion below, where we allow the parameter to affect both production costs and quality, we borrow the term "capability" from Sutton (2007).

direct complementarity between input quality and plant capability. In both variants, the output price–plant size elasticity and the input price–plant size elasticity are increasing in the scope for quality differentiation. If the scope for quality differentiation is sufficiently large, then in both cases larger, more capable plants purchase higher-quality inputs, pay more for inputs, sell higher-quality outputs, and charge higher prices.

In Section 5, we return to the data and show that key cross-sector predictions of the model are borne out empirically. In particular, using a measure of the scope for quality differentiation from Sutton (1998)—the R&D and advertising intensity of U.S. industries—we find that both the output price–plant size and the input price–plant size elasticities are greater in sectors with greater scope for quality differentiation. This result is robust to controlling for observable sector-level measures of horizontal differentiation.

In Section 6, we examine alternative explanations for the price–plant size patterns that rely on differences in market power across plants, instead of quality differences. To frame the empirical investigation, we discuss how extensions of the model of Foster, Haltiwanger and Syverson (2008), in which goods are symmetric but plants are subject to idiosyncratic demand shocks, can be reconciled with our two main stylized facts. Empirically, we show that the input price–plant size elasticity is smaller in sectors with more concentration of intermediate input suppliers, suggesting that market power of input suppliers is not responsible for the positive input price–plant size correlation. We find some support for the hypothesis that monopsony power of downstream purchasers may explain some of the input price dispersion, but the positive input price–plant size correlation exists even among producers with no monopsony power by our measures.

An important caveat is that we do not observe product quality in the data, and the evidence for quality differences is necessarily indirect. Nevertheless, in our view, the findings in Sections 5 and 6—that an off-the-shelf measure of the scope for quality differentiation, from arguably the leading existing work on the subject, predicts cross-sector variation in the price–plant size elasticities; and that observable measures of market power appear not to provide a complete explanation for the observed patterns—provide reasonably strong support for the view that the price–plant size correlations at least in part reflect differences in both input and output quality.

Perhaps the most important implication of this conclusion for our understanding of the industrial development process is that quality upgrading may require the upgrading of entire complexes of suppliers and downstream producers, not just of particular leading firms. While individual firms may be able to import high-quality inputs from abroad,[6] the existence of transport costs and the increasing prevalence of "just-in-time" production and other forms of local coordination (see *e.g.* Milgrom and Roberts, 1990) suggest that a lack of locally available high-quality inputs is likely to hinder the ability of even the most talented and knowledgeable entrepreneurs to upgrade.[7] Relatedly, the upgrading of downstream producers is likely to generate pressure on local suppliers to improve quality, much as foreign direct investment may lead to productivity improvements among domestic suppliers (Javorcik, 2004; Kugler, 2006).

In addition to the work cited above, this paper is related to several branches of literature. It is related to the literature using trade-flow data—in which unit values are available at a sector level (i.e. 6- or 10-digit trade categories)—to draw inferences about product quality. A standard argument in this literature is that if a sector in a country is able to export a large volume at a high price, then it must be producing high-quality goods (Hummels and Klenow, 2005; Hallak and

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[6] For evidence that imported inputs tend to be of higher quality than domestic ones, see Kugler and Verhoogen (2009) and Halpern, Koren and Szeidl (2009).

[7] Conversely, greater availability of high-quality inputs is likely to promote upgrading of outputs. This point is complementary to the finding of Goldberg *et al.* (2010) that the availability of greater *variety* of inputs stimulates production of new *varieties* of outputs.

Schott, 2011; Khandelwal, 2010). In using the positive correlation between plant size and output prices to argue for quality differences, the current paper makes a similar argument at the plant level. But it also strengthens the argument, in that we are able to use data on inputs to distinguish the quality story from explanations based on demand shocks and market power that would be observationally equivalent in data on outputs alone.[8]

This paper is also related to a small number of papers using product-level unit-value information in plant-level data sets. Several studies use information from the U.S. Census of Manufactures for a limited number of relatively homogeneous sectors for which unit values can be calculated on a consistent basis, but find *negative* correlations between plant size and output prices in those homogeneous sectors (Roberts and Supina, 1996, 2000; Syverson, 2007; Foster, Haltiwanger and Syverson, 2008).[9] The results below are consistent with this finding for the most homogeneous sectors but also indicate that the most homogeneous sectors are not representative of the Colombian manufacturing sector as a whole. An early draft of an independent paper by Hallak and Sivadasan (2009), which is focused on quality differences between exporters and non-exporters conditional on plant size, documented a positive correlation between output price and plant size in Indian data. An advantage of the Colombian data is that they contain information on the unit values of material inputs, which we argue is important for distinguishing the implications of quality models from competing explanations. Halpern and Koren (2007) use Hungarian firm data linked to unit values in trade-transactions data to analyse dispersion in prices of imported inputs, and find evidence that prices of imported inputs are higher for buyers with more market power in their output markets. An advantage of the Colombian data is that we observe unit values for domestic inputs and outputs in addition to internationally traded ones, and our analysis may be less subject to concerns about endogenous selection into international markets.[10]

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[8] Other leading contributions to the literature using unit values in trade-flow data include Rodrik (1994), Schott (2004), Hummels and Skiba (2004), Hallak (2006), and Choi, Hummels and Xiang (2009). A recent paper by Fajgelbaum, Grossman and Helpman (2009) provides a theoretical framework in which to analyse the role of income distribution and product quality in explaining sector-level trade flows and welfare gains from trade. Two recent papers relate unit values in trade-flow data to extensions of the Melitz (2003) model: Baldwin and Harrigan (2011) and Johnson (2010). The key finding of these papers is that exports to more distant markets have higher unit values on average. This fact is consistent with the hypothesis that more-productive firms produce higher-quality goods and charge higher prices, a hypothesis that is explicitly present in the earlier paper by Verhoogen (2004, 2008) and implicitly present in Melitz (2003), given a suitable redefinition of quality units — a redefinition alluded to in the original paper (Melitz, 2003, p. 1699). Appendix D spells out a "quality" version of the Melitz model suggested in the original paper, shows how it relates to the model we present in Section 4, and shows that the Baldwin and Harrigan (2011) model is isomorphic to it. The key difference between the quality version of the Melitz model and our model, as will become clear below, lies in the treatment of inputs.

[9] These papers do not report correlations between input prices and plant size. The only study we are aware of that explicitly considers the correlation between non-labour domestic input prices and plant size is Davis *et al.* (2006), which shows that electricity prices paid by manufacturing plants are *decreasing* in purchase volume. In other related work, Aw, Batra and Roberts (2001) investigate plant-level price differences between goods sold on the export and domestic market in the Taiwanese electronics sector. In Mexican data, Iacovone and Javorcik (2009) argue that plants raise output prices in preparation for exporting, which suggests that the quality-upgrading process highlighted by Verhoogen (2008) begins prior to entry into the export market. Crozet, Head and Mayer (2009) use price information and direct quality ratings on French wines to test the implications of a quality-sorting model of trade. The latter three papers do not focus on the unit values of material inputs.

[10] In other recent papers using firm-level trade-transactions data, Bastos and Silva (2010), Manova and Zhang (forthcoming), Martin (2010) and Görg, Halpern and Muraközy (2010) show that within firms there is a positive correlation between export price and gross domestic product per capita of the destination (in Portugal, China, France, and Hungary, respectively), consistent with the hypothesis of Verhoogen (2008) that individual firms sell higher-quality varieties to consumers more willing to pay for quality. Manova and Zhang (forthcoming) also show that Chinese firms that charge higher export prices pay more for their imported inputs, providing additional evidence for ideas previously advanced by (earlier versions of) this paper and Kugler and Verhoogen (2009).

Finally, this paper is related to an older literature in labour economics that documents a robust positive correlation between establishment size and wages, the "employer size-wage effect" (Brown and Medoff, 1989; Oi and Idson, 1999). To our knowledge, this paper is the first to present evidence from broadly representative data that the positive correlation generalizes to material inputs. The fact that the pattern holds for material inputs as well as labour tends to support the argument that the residual size–wage correlation (*i.e.* after controlling for observable worker characteristics) at least in part reflects differences in unobserved labour quality, and not solely institutions or frictions that are specific to the labour market.[11]

2. DATA AND ILLUSTRATIVE EXAMPLES

2.1. *Description of Data sets*

The data we use are from the *Encuesta Anual Manufacturera* (EAM) [Annual Manufacturing Survey], collected by the *Departamento Administrativo Nacional de Estadística* (DANE), the Colombian national statistical agency. The data set can be considered a census of manufacturing plants with 10 or more workers. In conjunction with this standard plant survey, DANE also collects information on the value and physical quantity of each output and input of each plant (valued at factory-gate prices), which is used to calculate national producer price indices. A unit value for each plant–product–year observation can then be calculated by dividing value (revenues or expenditures) by physical quantity.[12] The unit value represents an average price paid or charged by a plant over a year; hereafter we will (somewhat loosely) use the terms unit value and price interchangeably. Product-level information is available for the 1982–2005 period. Information on exports and imports, as well as employment and earnings of blue-collar and white-collar workers, is available on a consistent basis only for 1982–1994. We construct two separate plant-level unbalanced panels, a 1982–2005 panel and a 1982–1994 panel. We observe approximately 4500–5000 plants in each year, producing in and purchasing from approximately 4000 distinct eight-digit product codes.

A great advantage of these data is that DANE analysts have been extremely careful about maintaining consistent units of measurement within product categories. DANE instructs plants on which measurement units to use. If plants report using alternative units, then DANE follows up to request that the plant report using the correct units. If plants successfully make the case that doing so is impossible, DANE creates a new product category for the good using the new units. Thus, for example, there exist two eight-digit product categories corresponding to weed killers and herbicides: product code 35123067 is measured in kilograms, and 35123075 in litres. As a consequence, units of measurement are truly homogeneous within categories. This fact, the fact that inputs are included as well as outputs, and the fact that product-level information is available for the entire population of manufacturing plants with 10 or more workers over a 20+ year period make the Colombian data unique, arguably better suited to analysing our research question than any data set in any other country. See Appendix A1 (Supplementary Material) for

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[11] In related work, Abowd, Kramarz and Moreau (1996) relate an index of product prices at the firm level to estimates of worker fixed effects from employer–employee wage regressions, and interpret the results as evidence for a relationship between output quality and worker quality. The presence of market power (*e.g.* if firms with market power can pass on labour cost shocks to consumers in the form of higher prices) would undermine this interpretation, however. The current paper differs in that we have information on material-input prices and use the additional information in price–plant size correlations (and how they vary across sectors) to evaluate the extent to which price differences can be attributed to quality vs. market-power differences.

[12] The product-level information has been used previously by Eslava *et al.* (2004, 2006, 2009) in studies that focus on other issues.

variable definitions and Appendix A2 (Supplementary Material) for full details on the cleaning and processing of the datasets.

Appendix Table 1 (Supplementary Material) presents summary statistics for the plant-level data in our two panels, the 1982–1994 panel and the 1982–2005 panel. Consistent with patterns for the U.S. documented by Bernard and Jensen (1995, 1999), exporting plants are larger, in terms of both sales and employment, and pay higher wages; a minority of plants export and, conditional on exporting, plants derive a minority of their sales from the export market. Consistent with patterns for Taiwan (Aw and Batra, 1999) and Mexico (Verhoogen, 2008), exporting plants have a higher white-collar-to-blue-collar wage ratio. Exporting plants produce in a larger number of distinct output categories and purchase from a larger number of distinct input categories than non-exporters. Appendix Table A2 (Supplementary Material) reports summary statistics for the product-level information in the 1982–2005 panel, by ISIC four-digit industry. Columns 1–4 consider outputs and Columns 5–8 consider inputs. The table reports the number of distinct eight-digit products in each four-digit industry, the industry-average number of distinct plants selling or purchasing each product in each year, the within-product standard deviation and the within-product-year standard deviation of log real prices for each product.[13] We note that there is a fair amount of variation across sectors both in the number of selling or purchasing plants per product and in the extent of within-product price dispersion, the within-product dispersion largely reflects within-product-year dispersion, rather than variation in real prices across years.

2.2. *Examples*

To illustrate the richness of the data set and how we use it to draw inferences about product quality, consider two specific products, chosen because they have relatively simple production processes and because they plausibly differ in the scope for quality differentiation: hollow bricks and bar soap. Hollow bricks (*ladrillos huecos*) are a common building material in Colombia, similar to cinderblocks but made of clay rather than concrete; the scope for quality differentiation in hollow bricks is arguably quite limited. Bar soap for washing *(jabon en pasta para lavar),* commonly used to hand-wash laundry in Colombia, is typically produced by combining some form of animal fat (suet) with lye (sodium hydroxide) or similar chemical.[14] Bar soap is sold under a number of well-known brand names, and consumers are sensitive to quality differences, as impurities can affect the colour, odour and rancidity of the soap. The scope for quality differentiation in bar soap is arguably much higher than in hollow bricks.

Figure 1A plots the log price (log real unit value) of hollow bricks against log plant size (here measured by log employment), with both variables deviated from year means. Figure 1B plots the price paid by producers of hollow bricks for the main input into hollow bricks, common clay (*arcilla común*). The output price–plant size slope is negative but not significant, and the input price–plant size slope is also close to zero with a larger standard error. Figures 2(A)–(C) plot analogous figures for bar soap and its main inputs, refined and unrefined rendered suet. In all three cases, the slopes are significantly positive.

Our argument can be summarized succinctly using these examples. The next section estimates average price–plant size elasticities for the the manufacturing sector as a whole and shows that manufacturing overall is more like bar soap than like hollow bricks. Section 4 develops a

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[13] The within-product standard deviation and the within-product-year standard deviation are the standard deviations of the residuals from regressions of log real unit value on a full set of product dummies or a full set of product-year dummies, respectively.

[14] Face and hand soap *(jabon de tocador)* is classified under a separate product code.

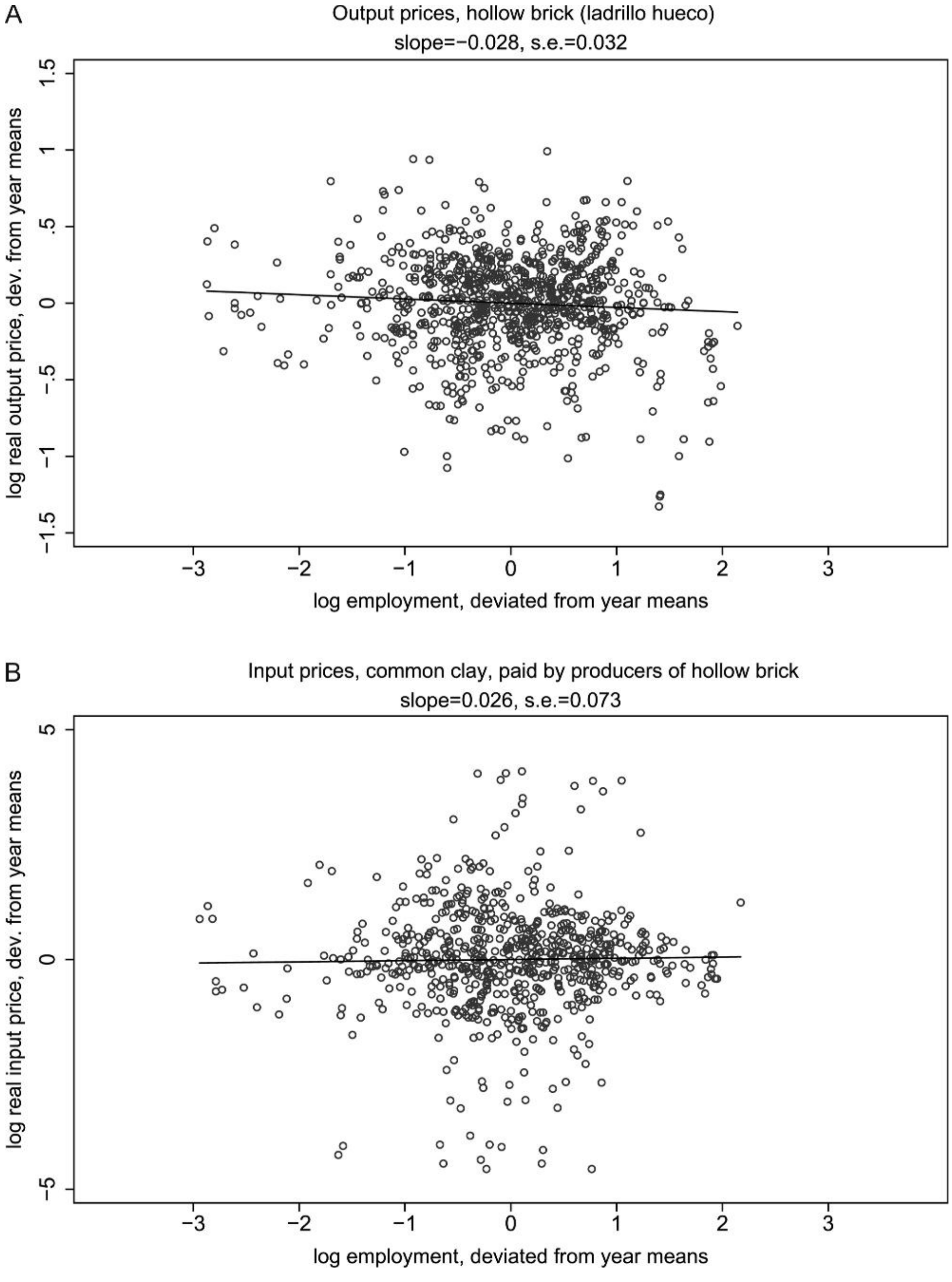

FIGURE 1: Illustrative example: Hollow Brick.

The figure uses the 1982–2005 panel. In top graph, each plotted point is log real output price charged by a producer of hollow brick vs. log employment, with both variables deviated from year means. In bottom graph, each plotted point is log real input price paid for common clay (the main input into hollow bricks) by a producer of hollow brick vs. log employment, again deviated from year means. Regression lines weight each plant-product-year observation equally. See Appendix A1 (Supplementary Material) for more detailed variable descriptions and Appendix A2 (Supplementary Material) for details of data processing

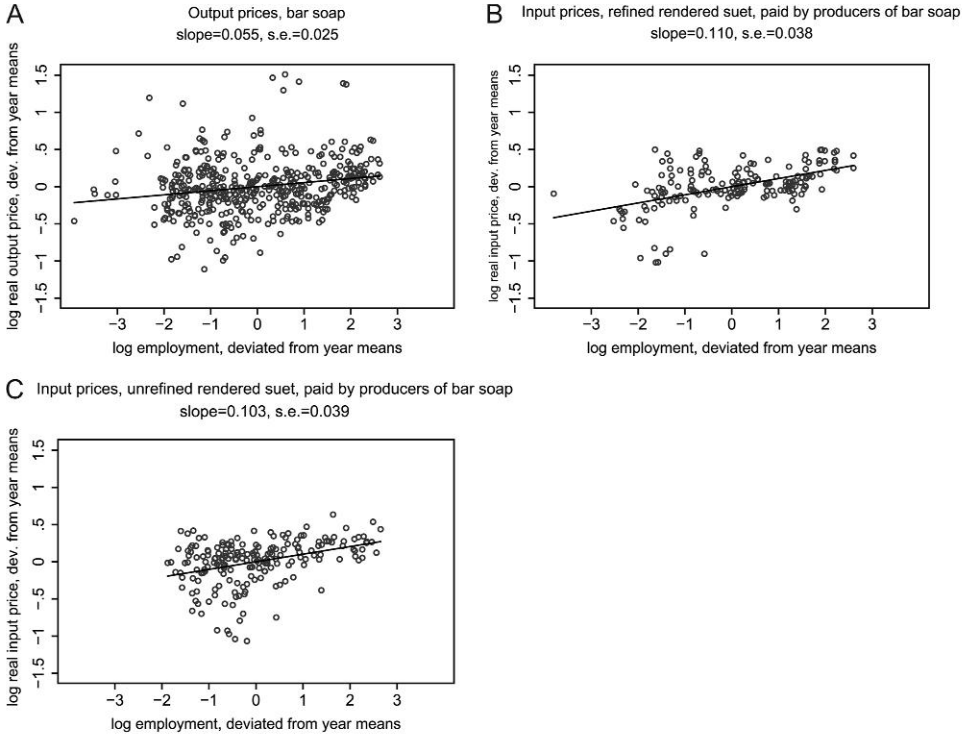

FIGURE 2: Illustrative example: Bar soap.

Figure uses the 1982–2005 panel. In 2A, each plotted point is log real output price charged by a producer of bar soap vs. log employment, with both variables deviated from year means. In 2B and 2C, each plotted point is log real input price paid for refined or unrefined rendered suet (the main inputs into bar soap for washing) by a producer of bar soap vs. log employment, again deviated from year means. Regression lines weight each plant-product-year observation equally. See Appendix A1 (Supplementary Material) for more detailed variable descriptions and Appendix A2 for details of data processing

model that accounts parsimoniously for both examples: if there is no scope for quality differentiation, then the model reduces to the standard interpretation of the Melitz model and predicts price–plant size elasticities similar to those we observe for hollow bricks in Figure 1; if the scope for quality differentiation is sufficiently large, then the model predicts elasticities similar to those we observe for bar soap in Figure 2. Using an off-the-shelf measure of the scope for quality differentiation, Section 5 shows that price–plant size elasticities are increasing in the scope for quality differentiation across broad manufacturing sectors. After considering alternative explanations in Section 6, we conclude that greater quality of both inputs and outputs in larger plants appears to be at least part of the explanation for the positive price–plant size elasticities we observe in the bar soap sector and on average for the manufacturing sector as a whole.

3. STYLIZED FACTS

This section documents average correlations between input prices, output prices, and plant size or export status. Our baseline econometric model is the following:

$$\ln p_{ijt} = X_{jt}\gamma + \psi_{it} + \delta_{rt} + \xi_k + \varepsilon_{ijt} \tag{1}$$

where i, j, k, and t index goods, plants, industries and years, respectively; $\ln p_{ijt}$ is log real unit value; X_{jt} is a measure of plant size, *e.g.* log total output or log employment; ψ_{it} is a product-year effect; δ_{rt} and ξ_k are region-year and industry effects, respectively;[15] and ε_{ijt} is a mean-zero disturbance. We run regressions separately for output prices and input prices. In some specifications, we estimate equation (1) with an indicator for export status or the export share of sales as the X_{jt} variable.

The coefficient of interest in these regressions is γ, which represents a price–plant size elasticity. In this section, we constrain γ to be the same for all industries; in this sense, it reflects an average elasticity across sectors. (We consider differences in these elasticities across industries below.) Note that the product-year effects, ψ_{it}, absorb all variation in prices of particular products that is common across plants. Intuitively, the coefficient γ in each regression is identified on the basis of a comparison of prices between plants of different sizes producing or consuming *the same product in the same year*, controlling for region-year and industry effects. In this way, we avoid the difficulty that we have no metric with which to compare unit values across products. It is worth emphasizing that the estimates of γ reflect *correlations*, not causal effects of plant size on prices. Indeed, we will hypothesize below that both plant size and prices are determined by unobserved heterogeneity in plant capability.

A natural measure of plant size is gross output; this is the standard measure of plant size used by DANE, measured as total sales plus net intra-firm transfers plus net change in inventories. Measurement error in gross output is a potential concern, however. To the extent that the measurement error is classical, it may simply attenuate coefficient estimates towards zero. But another form of bias is also likely: revenues represent output quantities times output prices and hence measurement error in prices may appear on both sides of the regression. This will generate a mechanical positive bias in the OLS estimate of γ.[16] To address these concerns, we use employment as an alternative measure of plant size. Employment has the advantage that measurement error is likely to be less severe and, importantly, uncorrelated with measurement error in values and quantities of outputs and inputs. We use log total employment as an instrument for log total output in a two-stage least-squares (2SLS) procedure; under the assumption that the measurement errors in gross output and total employment are uncorrelated, the 2SLS estimator will yield consistent estimates.

Observations at the plant-product-year level may not be independent either across products within plant-years or across years within plants. For this reason, we cluster errors at the plant level, allowing for arbitrary correlation across products and years within plants.[17] We report the number of distinct plants (*i.e.* clusters) as well as the number of plant-product-year observations used in each regression. Although respondents are instructed to value outputs and inputs at factory-gate prices, it is conceivable that some reported prices capture transport costs. The region-year effects, δ_{rt}, capture any systematic variation in transport costs or other region-specific influences on prices.

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[15] Note that the industry effects are not redundant even though product-year effects are included because plants are assigned to industries based on the relative importance of all the products they produce, and two plants producing the same product may belong to different industries depending on the their product mixes. For details, see Appendix A2 (Supplementary Material)

[16] A related concern is that if we regress unit values (revenues or expenditures divided by physical quantities) on physical quantities, then measurement error in quantities appears on both sides, generating a mechanical negative bias. This explains why our baseline regressions are not of prices on physical quantities; we consider such regressions in Appendix B4 (Supplementary Material).

[17] Note that this procedure is more conservative than clustering separately by plant-year across products and product-plant across years.

TABLE 1
Product prices vs. plant size

	OLS (1)	Reduced form (2)	2SLS (3)
A. Dependent variable: log real output unit value			
Log total output	0·021***		0·025***
	(0·005)		(0·006)
Log employment		0·026***	
		(0·007)	
Product-year effects	Y	Y	Y
Industry effects	Y	Y	Y
Region-year effects	Y	Y	Y
R^2	0·90	0·90	
N (obs.)	413,789	413,789	413,789
N (plants)	13,582	13,582	13,582
B. Dependent variable: log real input unit value			
Log total output	0·015***		0·011***
	(0·002)		(0·003)
Log employment		0·012***	
		(0·003)	
Product-year effects	Y	Y	Y
Industry effects	Y	Y	Y
Region-year effects	Y	Y	Y
R^2	0·78	0·78	
N (obs.)	1,338,921	1,338,921	1,338,921
N (plants)	13,582	13582	13,582

Notes: The table uses the 1982–2005 panel. Total output is total value of production, defined as sales plus net transfers plus net change in inventories. In Column 3, log employment is instrument for log total output; the coefficient on log employment, its robust standard error and the R^2 in the first stage are 1·058, 0·011, and 0·733, in Panel A and 1·082, 0·10 and 0·782 in Panel B, respectively. Product-year and industry effects are not perfectly collinear because industry is defined as the industry category with the greatest share of plant sales, and two plants producing the same product may be in different industries. Errors clustered at plant level. *N* (plants) reports number of clusters (*i.e.* distinct plants that appear in any year). Robust standard errors are given in parentheses. *10% level, **5% level, ***1% level. See Appendix A1 for more detailed variable descriptions and Appendix A2 for details of data processing.

Panel A of Table 1 presents estimates of equation (1) with log real output price as the dependent variable. Columns 1 and 2 use log total output and log employment, respectively, as the measures of plant size. Column 3 reports the 2SLS estimate of the coefficient on log total output using log employment as the instrument for log total output. The 2SLS estimate is slightly larger than the OLS estimate, consistent with the hypothesis that measurement error in gross output is generating attenuation bias. The coefficient estimates are highly statistically significant and indicate that output prices are positively correlated with plant size on average. Column 2 indicates that 10% greater employment is associated with approximately 0·26% higher output prices.

Panel B of Table 1 presents analogous regressions with the log real input price as the dependent variable. In moving from Column 1 to Columns 2 and 3, the coefficient on plant size falls slightly. This may reflect non-classical measurement error.[18] But again, the important message

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[18] For instance, if measurement error in revenues is positively correlated with error in expenditures, then the mechanical positive relationship between input prices and expenditures due to measurement error in expenditures may also generate a positive bias in the relationship between input prices and revenues.

TABLE 2

Product prices vs. export measures, plant size

	(1)	(2)	(3)	(4)	(5)
A. Dependent variable: log real output price					
Log employment	0·025***			0·009	0·020**
	(0·008)			(0·008)	(0·008)
Exporter		0·114***		0·104***	
		(0·022)		(0·023)	
Export share			0·288**		0·251*
			(0·137)		(0·142)
Product-year effects	Y	Y	Y	Y	Y
Industry effects	Y	Y	Y	Y	Y
Region-year effects	Y	Y	Y	Y	Y
R^2	0·90	0·90	0·90	0·90	0·90
N (obs.)	216,155	216,155	216,155	216,155	216,155
N (plants)	10,106	10,106	10,106	10,106	10,106
B. Dependent variable: log real input price					
Log employment	0·013***			0·008**	0·013***
	(0·004)			(0·004)	(0·004)
Exporter		0·037***		0·028***	
		(0·009)		(0·009)	
Export share			0·021		−0·002
			(0·027)		(0·027)
Product-year effects	Y	Y	Y	Y	Y
Industry effects	Y	Y	Y	Y	Y
Region-year effects	Y	Y	Y	Y	Y
R^2	0·80	0·80	0·80	0·80	0·80
N (obs.)	684,746	684,746	684,746	684,746	684,746
N (plants)	10,106	10,106	10,106	10,106	10,106

Notes: The table uses the 1982–1994 panel since export status is reported on a consistent basis only for those years. Exporter equals 1 if plant has exports >0, and 0 otherwise. Export share is fraction of total sales derived from exports. Product-year and industry effects are not perfectly collinear because industry is defined as the industry category with the greatest share of plant sales, and two plants producing the same product may be in different industries. Errors clustered at plant level. *N* (plants) reports number of clusters (*i.e.* distinct plants that appear in any year). Robust standard errors are given in parentheses. *10% level, **5% level, ***1% level. See Appendix A1 for more detailed variable descriptions and Appendix A2 for details of data processing.

is that input prices are positively correlated with plant size on average. The estimates suggest that 10% greater plant size is associated with approximately 0·11–0·12% higher prices paid for material inputs.[19]

In Table 2, we show that similar patterns hold for the relationships between prices and export status. The table uses the 1982–1994 panel, the panel in which export status is observed. To simplify the presentation of results, hereafter we focus on the reduced-form regressions with log employment as the key covariate; 2SLS estimates using log employment as an instrument for log total output (available from the authors) are similar. Results for output prices are in Panel A, and results for input prices in Panel B. For comparison purposes, Column 1 of each panel presents a regression with log employment as the key covariate, comparable to Column 2 of

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[19] 19. Note that the output price–plant size relationship estimated in Panel A is steeper than the input price–plant size relationship in Panel B, suggesting that profitability may be increasing in plant size as well. This difference is not robust, however. (See footnote 10 in Appendix B1.)

Table 1; the results are similar to those for the longer 1982–2005 panel. The results in Column 2 of Table 2 indicate that both output and input prices are higher among exporters.[20] On average, exporters (*i.e.* plants with non-zero exports) have approximately 11% greater output prices and 3.7% greater input prices than non-exporters. In Column 3, with the export share of revenues as the key covariate, the coefficient is significant and positive for output prices and not significant but of the expected sign for input prices. Caution is warranted in interpreting the results in Columns 4 and 5 since in Melitz-type models (including the one we propose in the next section) both employment and export status reflect a single underlying capability parameter and the source of variation that allows them to be separately identified is unclear. Subject to that caveat, the results in Column 4 indicate that being an exporter is associated with both higher output prices and higher input prices, even conditional on plant size.[21] Again, the result for the relationship of export share to input prices (Column 5) is less robust.

In Appendix B, we present a number of robustness checks of these basic findings. First, we show that the results are robust to using a two-step method to estimate the price–plant size elasticities, first regressing prices on plant-year effects (and other covariates) and then regressing the estimated coefficients from the plant-year effects on plant size (and other covariates). This method also provides a straightforward way to compare plant-average output prices to plant-average input prices; we find, unsurprisingly, that there is a positive relationship.[22] Second, we estimate equation (1) using non-exporting plants only and find similar results; this suggests that the results are not being driven solely by plants selling higher-quality goods to export markets. Third, we confirm that the familiar employer size-wage effect holds in our data, separately for white-collar and blue-collar workers. Finally, we examine the relationship between prices and physical quantities at the product level. An important pitfall in interpreting such regressions is that, since prices are calculated as value divided by physical quantity, any measurement error in physical quantity will generate a spurious, mechanical negative correlation between prices and physical quantities, a problem pointed out in the context of household surveys by Deaton (1988). Using employment as an instrument for physical quantity to address this issue, we find results similar to those reported above.

4. A MODEL OF ENDOGENOUS INPUT AND OUTPUT QUALITY CHOICE

In this section, we show that the empirical patterns above are consistent with a parsimonious extension of the Melitz (2003) framework to include endogenous choice of input and output quality. We develop two variants of the model, corresponding to different specifications of the production function for product quality. The two variants carry similar observable implications in our data, and we do not seek to discriminate between them empirically in this paper. The important theoretical point is that both variants yield an equilibrium in which (provided there is positive scope for quality differentiation) more capable entrepreneurs purchase higher-quality inputs to produce higher-quality outputs.

An important caveat is that the model uses a demand system and other functional form assumptions that are special in a number of ways. We have chosen a Melitz-type framework for its

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[20] The latter fact is a generalization of the finding that exporters in the U.S. pay higher wages than non-exporters, established by Bernard and Jensen (1995, 1999) for the U.S. and confirmed in many other countries. See Schank, Schnabel and Wagner (2007) for a review.

[21] For a theoretical framework that can provide a coherent account for this pattern, see Hallak and Sivadasan (2009).

[22] There are many models in which a shock to input prices would be passed on to consumers in the form of higher output prices. What is perhaps more surprising is that both are positively correlated with plant size.

tractability in analysing the behaviour of large numbers of heterogeneous firms in general equilibrium and for its comparability to the recent trade literature,[23] but we acknowledge that similar observable implications could be derived from models of product quality choice under different market structures (*e.g.* perfect competition or the vertical differentiation model of Shaked and Sutton (1982)) and in the absence of heterogeneity in capability among entrepreneurs.[24]

4.1. *Basic Set-up*

There are two symmetric countries and in each country two sectors, a monopolistically competitive final-good sector and a perfectly competitive, constant-returns-to-scale intermediate-input sector. Both final goods and inputs may have quality differences, in manners that will be made clear below.

In each country, a representative consumer has the following standard asymmetric constant-elasticity-of-substitution utility function over final goods:

$$U = \left[\int_{\omega\in\Omega} (q(\omega)x(\omega))^{\frac{\sigma-1}{\sigma}}\, d\omega\right]^{\frac{\sigma}{\sigma-1}} \tag{2}$$

where ω indexes varieties in the final-good sector; $\mathbf{\Omega}$ represents the set of all differentiated varieties available in the market (produced in either country); σ is a parameter capturing the elasticity of substitution between varieties, where we make the standard assumption that $\sigma > 1$; $x(\omega)$ is the quantity consumed; and $q(\omega)$ is the quality of variety ω, assumed to be observable to all. Here, output quality, $q(\omega)$, can be interpreted as any product attribute that the representative consumer values and that is chosen by firms.[25]

Consumer optimization yields the following demand for a particular variety, ω, in each country:

$$x(\omega) = Xq(\omega)^{\sigma-1}\left(\frac{p_O(\omega)}{P}\right)^{-\sigma} \tag{3}$$

where $p_O(\omega)$ is the price (the "output price") of variety ω, P is an aggregate quality-adjusted price index, and X is a quality-adjusted consumption aggregate of the varieties available on the

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[23] It is worth emphasizing that the Melitz framework is not ideally suited to analysing market concentration, because it is difficult to reconcile with the fact that in many industries the number of market players remains fixed even as the market grows large (see Sutton (1991, pp. 70–71)), but explaining concentration is not our goal here.

[24] In a model of perfect competition, if (1) producing high-quality goods requires high-quality inputs and (2) the technology for producing high-quality goods for some reason requires a larger size than the technology for producing low-quality goods, then one would expect to see positive correlations of output prices, input prices, and plant size. It is not obvious why producing higher quality would require a larger scale, but it is certainly a possibility. In Shaked and Sutton (1982), firms are *ex ante* homogeneous but endogenously locate at different locations in a vertical quality space, with some firms paying greater fixed quality costs and offering higher-quality goods. Shaked and Sutton (1982, lemma 3) show in their context that high-quality firms have higher revenues than low-quality ones. If, in addition, high-quality production requires high-quality inputs (as in Sutton (1991, section 3.5)) then one can generate observable implications similar to those of our model. In either the perfect-competition or vertical-quality-space models, however, the price dispersion in both outputs and inputs would reflect quality differences, which we also view as the key point of this paper.

[25] Note that this notion of quality is distinct from more general notions of "product appeal" or "demand shocks" that respond to changes in consumer valuations even conditional on a given set of production choices by firms. Treating quality as single-dimensional and assuming a representative consumer are clearly stark simplifications, but Anderson, de Palma and Thisse (1992) show that the demand patterns generated by such a representative consumer can also be generated by a model with many individual consumers making discrete choices. In this view, quality can be interpreted as a component of product attributes that is valued by all consumers, where the residual component of consumers' heterogeneous valuations has mean zero.

market.[26] Note that demand is increasing in the quality and decreasing in the price of a particular variety.

Like Melitz (2003), we begin with an inelastic labour supply L (measured in labour-hours) with the hourly wage normalized to one. But we add the intermediate-input sector, which transforms homogeneous labour-hours into intermediate inputs of different qualities. In the intermediate-input sector, the production function is simply:

$$F_{\mathrm{I}}(\ell, c) = \frac{\ell}{c}, \tag{5}$$

where c is the quality of the intermediate input produced and ℓ is the number of labour-hours used. In other words, producing one unit of an intermediate input of quality c requires c labour-hours and, given the wage normalization, entails cost c. Let $p_{\mathrm{I}}(c)$ be the price of an intermediate input of quality c. Final-good producers are assumed to be price-takers in intermediate-input markets, and all face the same input price–input quality schedule $p_{\mathrm{I}}(c)$. It will turn out that in equilibrium the price of each intermediate input equals the marginal cost of producing the input: $p_{\mathrm{I}}(c) = c$.

The simplest interpretation of the model, which we adopt here in the interest of clarity, is that the intermediate-input sector produces only material inputs and workers are homogeneous. But a valid alternative interpretation of the model is that the intermediate-input sector is an education sector, and that c unskilled labour-hours are required to "produce" one labour-hour of skill c.[27] In either interpretation, the key point is simply that, from the perspective of final good producers, there is a linear relationship between the quality of an intermediate input and the price of that input.[28]

As in Melitz (2003), to enter the final-good sector, entrants must pay an investment cost, f_e (measured in labour-hours) in order to receive a capability draw, λ.[29] We assume that capability is drawn from a Pareto distribution with c.d.f. $G(\lambda) = 1 - \left(\frac{\lambda_m}{\lambda}\right)^k$, with $0 < \lambda_m \leq \lambda$.[30] There is an exogenous probability of exit, δ, in each period.[31] There is a fixed cost of production, f, and an additional fixed cost of exporting, $f_x > f$, in each period. In the interests of simplicity, we assume that there are no variable costs of trade. Since there is no cost of differentiation, each plant in final-good sector produces a distinct good and λ can be used to index both plants and varieties.

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[26] Specifically, $X = U$ from equation (2) and

$$P \equiv \left[\int_{\omega\in\Omega} \left(\frac{p_O(\omega)}{q(\omega)}\right)^{1-\sigma} d\omega\right]^{\frac{1}{1-\sigma}} \tag{4}$$

[27] Although we abstract from life-cycle considerations, one could also think of c as the amortized per-period portion of an initial investment in education at the beginning of a worker's career.

[28] In the interests of tractability, we have assumed that input quality is observable to all, but similar conclusions would hold in models in which downstream producers imperfectly observe input quality, as long as there were a positive relationship between input price and input quality, as for instance in the reputation model of Shapiro (1983).

[29] As mentioned above, Melitz (2003) refers to this parameter as productivity. To avoid confusion below, where we allow the parameter to affect both the cost of production and product quality, we borrow the term "capability" from Sutton (2007).

[30] Below we will impose a lower bound on the shape parameter, k, to ensure that the distribution of capability draws has a finite variance. Helpman, Melitz and Yeaple (2004) impose a similar restriction.

[31] As in Melitz (2003), we will focus on a steady-state equilibrium in which new entrants replace the exiters, and the distribution of plant capabilities remains constant over time.

Production in the final-good sector is described by two functions, one describing the production of physical units of output and the other describing the production of quality. The production of physical units is assumed to be:

$$F(n) = n\lambda^a \tag{6}$$

where n is the number of units of inputs used and a is a parameter reflecting the extent to which capability lowers unit costs, with $a > 0$. This function implies that $\frac{1}{\lambda^a}$ units of inputs are used for each physical unit of output, and hence the marginal cost of each unit of output is $\frac{p_I(c)}{\lambda^a}$. Below we will consider two variants of the production function for product quality, in which we allow product quality, q, to depend on different combinations of the plant's capability draw, λ, input quality, c, and a fixed investment in quality, f_q.[32]

Plants in the final-good sector optimize over the choices of input quality, c, fixed quality investment, f_q, output price, p_O, and which markets to enter. Let $Z = 1$ if the plant enters the export market, and 0 otherwise. The profit function for each final-good producer is then:

$$\pi(p_\mathrm{O}, c, f_\mathrm{q}, Z; \lambda) = \left(p_\mathrm{O} - \frac{p_\mathrm{I}(c)}{\lambda^a}\right)x - f_\mathrm{q} - f + Z\left[\left(p_\mathrm{O} - \frac{p_\mathrm{I}(c)}{\lambda^a}\right)x - f_x\right], \tag{7}$$

where demand, x, is given by equation (3) and depends on quality, q, and output price, p_O. Each plant in the continuum of plants is small relative to the size of the market and ignores the effects of its decisions on the aggregates X and P. Note that the symmetry of countries implies that, conditional on a choice of fixed quality costs, f_q, the optimal choices of c and p_O for exporters will be the same for both markets.

Characterizing the equilibrium of the model with a general production function for quality quickly becomes intractable. Instead, we consider separately two special cases, one in which output quality reflects a complementarity between input quality and plant capability but does not require a fixed quality investment, and one in which output quality reflects input quality and the fixed quality investment but there is no direct complementarity between plant capability and input quality. We first describe the equilibrium in each variant, and then turn to a discussion of the key common implications.

4.2. *Variant 1: complementarity between input quality and plant capability*

In this variant, we assume that plant capability, λ, and input quality, c, are complements in generating output quality and that upgrading does not require fixed costs.[33] In particular, we assume:

$$q = \left[\frac{1}{2}\left(\lambda^b\right)^\theta + \frac{1}{2}\left(c^2\right)^\theta\right]^{\frac{1}{\theta}} \tag{8}$$

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[32] Given that we focus on a steady-state equilibrium, it does not matter whether we think of the fixed quality investment as a fixed cost or the per-period amortized cost of an initial sunk investment. Similarly, f_x can be interpreted as a per-period fixed cost or as the amortized per-period portion of a single sunk cost paid when first entering the export market. See Melitz (2003, p. 1708).

[33] Complementarities among inputs in production have been emphasized by Sattinger (1979), Milgrom and Roberts (1990), Kremer (1993), Redding (1996), Grossman and Maggi (2000), Acemoglu, Antràs and Helpman (2007), and Jones (2009) among others, although these papers do not focus on complementarities between inputs and plant capability draws in the sense of Melitz (2003), as we do here. This variant is related to Verhoogen (2008), which hypothesizes a complementarity between plant productivity and *labour* quality in producing output quality, in a partial equilibrium framework with non-homothetic demand. The theoretical contribution of this variant, beyond generalizing the earlier hypothesis to material inputs, is to embed it in a general-equilibrium framework. Empirically, a contribution of the current paper is to show that the cross-sectional price predictions of the earlier paper hold "out of sample"—in Colombia rather than Mexico, and for output and input prices as well as wages.

The choices of multiplicative factor $\frac{1}{2}$ and the quadratic form in c are convenient but not crucial.[34] The parameter θ reflects the degree of complementarity between capability and input quality; as θ becomes more negative, the degree of complementarity increases. We impose the assumption that $\theta < 0$; this ensures that $q(\cdot,\cdot)$ is log-supermodular in λ and c.[35] Intuitively, we are assuming that the marginal increase in output quality for a given increase in input quality is greater for more capable entrepreneurs.[36] This rules out the possibility that plant capability and input quality are substitutes, for instance, because more capable entrepreneurs are particularly able to compensate for deficiencies in input quality.[37]

The parameter b in equation (8) reflects the scope for quality differentiation, which can be thought of as reflecting the availability of technology for translating higher plant capability into improved product quality. Although not explicitly in the model, one could also think of b as capturing the willingness of consumers to pay for product quality. The key point is that a higher b gives more capable plants a relatively greater incentive to produce high-quality outputs.[38] We assume $b \geq 0$.

As mentioned above, profit maximization and free entry in the intermediate-input sector imply that $p_{\mathrm{I}}(c) = c$ for all levels of input quality produced in equilibrium.[39] In the final-good sector, the first-order conditions for the plant's maximization problem in equation (7) imply the following:

$$c^*(\lambda) = p_I^*(\lambda) = \lambda^{\frac{b}{2}} \tag{9a}$$

$$q^*(\lambda) = \lambda^b \tag{9b}$$

$$p_{\mathrm{O}}^*(\lambda) = \left(\frac{\sigma}{\sigma-1}\right)\frac{p_{\mathrm{I}}^*(\lambda)}{\lambda^a} = \left(\frac{\sigma}{\sigma-1}\right)(\lambda)^{\frac{b}{2}-a} \tag{9c}$$

$$r^*(\lambda) = (1+Z)\left(\frac{\sigma-1}{\sigma}\right)^{\sigma-1} XP^{\sigma}(\lambda)^{\eta} \tag{9d}$$

where $c^*(\lambda)$ and $p_{\mathrm{I}}^*(\lambda)$ represent the input quality chosen and input price paid by plant λ in equilibrium, $p_{\mathrm{O}}^*(\lambda)$ is optimal output price, $q^*(\lambda)$ is optimal quality, $r^*(\lambda)$ is total revenues (i.e.

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[34] If the quality production function in equation (8) were instead:

$$q = \left[\mu\left(\lambda^b\right)^{\theta} + (1-\mu)\left(c^{\gamma}\right)^{\theta}\right]^{\frac{1}{\theta}}$$

then the conditions $0 < \mu < 1$ and $\gamma > 1$ would be sufficient.

[35] For a very useful discussion of log-supermodularity in a trade context, see Costinot (2009).

[36] Following the circulation of the first version of this paper (Kugler and Verhoogen , 2008a), Mandel (2010) and Johnson (2010) have presented models similar to this variant in which more capable firms are assumed to have a comparative advantage in producing higher-quality goods.

[37] We acknowledge that there may exist industries in which capability and input quality are substitutes rather than complements. In such a case, our model would not have an interior solution, the predictions below for the observable relationships between plant size and prices would not hold, and this variant of our model would not offer an explanation for the empirical patterns documented in Section 3.

[38] The parameter b corresponds loosely to the "escalation parameter" in Sutton (1998) (the correspondence is more direct in the second variant of our model—see footnote 43) and to the length of quality ladders in Khandelwal (2010).

[39] In the input sector, the profit of producing a unit of quality c is given by $\pi_I(\ell, c) = p_{\mathrm{I}}(c)F_{\mathrm{I}}(\ell, c) - \ell = \frac{p_{\mathrm{I}}(c)\ell}{c} - \ell$. Free entry in the input sector implies $p_{\mathrm{I}}(c) = c$. The number of units of each quality produced is determined by demand from the final-good sector.

from both markets for exporters), and $\eta \equiv (\sigma - 1)\left(\frac{b}{2} + a\right) > 0$.[40] Since fixed quality costs are by assumption ineffective in raising quality, plants simply set $f_{\mathrm{q}} = 0$.

The solution for the remaining endogenous variables is similar to Melitz (2003), and the details have been relegated to Appendix C1 (Supplementary Material) for that reason. To summarize briefly, three conditions—a zero-profit condition for remaining in the domestic market, a zero-profit condition for entering the export market, and a free-entry condition that the *ex ante* expected present discounted value of paying the investment cost to receive a capability draw is zero—pin down the cut-off values for remaining in the domestic market, λ^*, and entering the export market, λ_x^*. Since $f_x > f$ by assumption, the cut-off for entering the export market is to the right of the cut-off for remaining in the domestic market: $\lambda^* < \lambda_x^*$. Total revenues in the final-good sector are equal to total labour income; this pins down the mass of final goods (and final-good producers) in equilibrium.

4.3. *Variant 2: fixed costs of upgrading*

In this variant, we drop the complementarity between plant capability and input quality and focus on the fixed costs of upgrading quality, more directly in the spirit of Sutton (1991, 1998).[41] We also assume that producing high-quality outputs requires high-quality inputs. The roles of fixed costs and input quality can be compactly expressed by a Leontief-type production function for quality:

$$q = \min\left({f_{\mathrm{q}}}^{\alpha}, c^2\right) \tag{10}$$

Again, the quadratic term for c is convenient but not crucial.[42] In this case, the scope for quality differentiation is represented by the parameter α, which reflects the extent to which quality increases with an increase in fixed quality investment. This parameter corresponds closely to the "escalation parameter" characterizing the scope for quality differentiation in Sutton (1998).[43] Following Sutton (1991, 1998), one could think of this parameter as characterizing the effectiveness of R&D spending in improving the technical dimensions of quality or the effectiveness of advertising expenditures in raising the perceived quality of the firm's output. In order to ensure an interior solution in the choice of quality, we must assume that the extent to which quality improves with fixed costs is bounded from above: $\alpha < \frac{2}{\sigma-1}$. We also assume $\alpha \geq 0$.

As above, free entry in the intermediate input sector implies $p_{\mathrm{I}}(c) = c$ for all levels of input quality. The first-order conditions for each final-good producer's optimization problem yield the following:

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[40] The fact that θ drops out of these expressions is a consequence of the choices of the multiplicative factor $\frac{1}{2}$ and exponent 2 in equation (8). In the more general case discussed in footnote 34, $c^*(\lambda)$ and hence $p_{\mathrm{O}}^*(\lambda)$ would depend on θ.

[41] This variant of the model (which did not appear in early versions of the paper (Kugler and Verhoogen (2008a, 2008b)) is similar to the earlier model of Hallak and Sivadasan (2009), which also considers fixed costs of raising quality. This variant differs in that it allows for only one dimension of heterogeneity, places more emphasis on the role of inputs, and is able to characterize the general-equilibrium solution in closed form. If we think of upgrading as an increase in the reliability of products, *i.e.* a mean-preserving decrease in the variance of quality, then the model is also in the spirit of Bardhan and Kletzer (1984), which hypothesizes scale economies in expenditures to reduce the variance of quality. This variant is also related to several recent papers that posit a fixed cost of innovation and/or increasing productivity in a Melitz-type context, without focusing on product quality: Bustos (2011), Costantini and Melitz (2008), Lileeva and Trefler (2010), and Aw, Roberts and Xu (2011).

[42] The key requirement is that the exponent on c be greater than one.

[43] To be precise, our α corresponds to $1/\beta$ in Sutton (1998, Chapter 3) , where β is defined as the elasticity of fixed outlays to the resulting quality. There is a one-to-one mapping between Sutton's $1/\beta$ and his escalation parameter (also denoted by α) in sectors with a single "technological trajectory" (*i.e.* no additional submarkets to which demand spillovers or economies of scope in R&D extend). Our α also plays a role similar to the parameter capturing the effectiveness of R&D investment, θ, in Leahy and Neary (1997).

$$c^*(\lambda) = p_I^*(\lambda) = (\alpha\Phi)^{\frac{\alpha}{2\zeta}}\,\lambda^{\frac{a\alpha(\sigma-1)}{2\zeta}} \tag{11a}$$

$$f_{\mathrm{q}}^*(\lambda) = (\alpha\Phi)^{\frac{1}{\zeta}}\,\lambda^{\frac{a(\sigma-1)}{\zeta}} \tag{11b}$$

$$q^*(\lambda) = (\alpha\Phi)^{\frac{\alpha}{\zeta}}\,\lambda^{\frac{a\alpha(\sigma-1)}{\zeta}} \tag{11c}$$

$$p_{\mathrm{O}}^*(\lambda) = \left(\frac{\sigma}{\sigma-1}\right)\frac{p_{\mathrm{I}}^*(\lambda)}{\lambda^a} = \left(\frac{\sigma}{\sigma-1}\right)(\alpha\Phi)^{\frac{\alpha}{2\zeta}}\,\lambda^{a\left(\frac{\alpha(\sigma-1)}{2\zeta}-1\right)} \tag{11d}$$

$$r^*(\lambda) = \frac{2}{\alpha}\left(\frac{\sigma}{\sigma-1}\right)(\alpha\Phi)^{\frac{1}{\zeta}}\,\lambda^{\frac{a(\sigma-1)}{\zeta}}, \tag{11e}$$

where $\Phi \equiv \frac{1}{2}\left(\frac{\sigma-1}{\sigma}\right)^{\sigma} XP^{\sigma}(1+Z)$ and $\zeta \equiv 1-\frac{\alpha}{2}(\sigma-1)$. Under the assumptions on α above, we have that $0 < \zeta \leq 1$. Appendix C2 (Supplementary Material) presents the solution for the remaining endogenous variables, which is again similar to Melitz (2003). In this variant, we impose a slightly stronger assumption on the relative magnitudes of f_x and f, in particular $f_x > \left(2^{1/\zeta}-1\right) f$, to ensure that the cut-off for entering the export market is to the right of the cut-off for remaining in the domestic market: $\lambda^* < \lambda_x^*$.[44]

4.4. *Discussion*

The two variants have a number of similar properties. First, both nest the standard Melitz (2003) model.[45] In both cases, setting the parameter reflecting the scope for quality differentiation equal to zero ($b=0$ in the first case, $\alpha=0$ in the second) yields an equilibrium in which all plants choose the same input quality ($c=1$), pay the same input price ($p_{\mathrm{I}}=1$), and produce the same output quality ($q=1$), and in which more capable plants have lower costs and charge lower output prices, as in the standard interpretation of Melitz (2003). Melitz (2003, p. 1699) points out that his model is consistent with quality differentiation given a suitable choice of quality units. In particular, if we interpret $p_{\mathrm{O}}^*(\lambda)$ as reflecting price in quality-adjusted units rather than physical units, the model can generate a zero or positive correlation between observed output price in physical units and plant size. (Appendix D (Supplementary Material) spells out this argument in detail.) But the model cannot generate a positive correlation between input prices and plant size. This is the key difference between either variant of our model and the quality interpretation of the Melitz model.

Second, in both cases, as long as there is positive scope for quality differentiation, plants with higher capability draws use higher-quality inputs and produce higher-quality outputs. That is, if $b > 0$ in the first variant or $\alpha > 0$ in the second, then $c^*(\lambda)$, $p_{\mathrm{I}}^*(\lambda)$ and $q^*(\lambda)$ are increasing in λ. In the first case, the matching is driven by the fact that more capable entrepreneurs have a comparative advantage in using higher-quality inputs. In the second case, it is driven by the fact that more capable plants produce at a larger scale and can spread the fixed quality costs over more units than less capable plants. As a consequence they pay higher fixed costs and use better quality inputs.

Third, in both cases, the output price is simply $\frac{\sigma}{\sigma-1}$ times marginal cost (as is standard in models with Dixit–Stiglitz demand specifications) but there are offsetting influences on marginal cost. On the one hand, more capable plants have lower costs for a given level of quality. On the

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[44] If $f_x \leq (2^{1/\zeta}-1)f$, then all plants that enter the domestic market also enter the export market.

[45] To be precise, both nest a special case of the Melitz model with a Pareto distribution of productivity draws and zero transport costs.

other hand, more capable firms choose higher-quality inputs, which raises marginal cost. In both cases, the relationship between output price and plant capability depends on the magnitude of the parameter capturing the scope for quality differentiation. In the first variant, output prices are increasing in λ if and only if $b > 2a$; in the second variant, output prices are increasing in λ if and only if $\alpha > \frac{1}{\sigma-1}$.

Fourth, in both cases, plant revenues are unambiguously increasing in λ. While more capable plants may charge higher prices, which all else equal would tend to reduce demand, this is always more than offset by an increase in sales due to higher quality.[46] There is a discontinuous jump in revenues at the cut-off for entry into the export market, λ_x^*, but the relationship between revenues and capability remains monotonic.

Fifth, the two cases carry similar predictions for the correlations between the key observable variables available in the Colombian data and how they would be expected to vary across sectors. Capability, λ, input quality, $c^*(\lambda)$, output quality, $q^*(\lambda)$, and fixed quality costs, $f_{\mathrm{q}}^*(\lambda)$, are unobservable, but plant size (*i.e.* revenues), $r^*(\lambda)$, input price, $p_{\mathrm{I}}^*(\lambda)$, and output price, $p_{\mathrm{O}}^*(\lambda)$, are observed. Formally, in the first variant, equations (9a–9d) imply

$$\frac{d\ln p_{\mathrm{I}}^*}{d\ln r^*} = \frac{b}{2\eta} \tag{12a}$$

$$\frac{d\ln p_{\mathrm{O}}^*}{d\ln r^*} = \frac{1}{\eta}\left(\frac{b}{2} - a\right) \tag{12b}$$

In the second variant, equations (11a–11e) imply

$$\frac{d\ln p_I^*}{d\ln r^*} = \frac{\alpha}{2} \tag{13a}$$

$$\frac{d\ln p_O^*}{d\ln r^*} = \alpha - \frac{1}{\sigma-1} \tag{13b}$$

These elasticities hold for all values of $r^*(\lambda)$ except at the cut-off for entry into the export market, where $r^*(\lambda)$ is discontinuous.[47] If the scope for quality differentiation is sufficiently large (*i.e.* if $b > 2a$ or $\alpha > \frac{1}{\sigma-1}$), then both variants predict positive relationships between output

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[46] It is important to note that this result does not necessarily hold for physical units of output. In each variant, there are parameter values satisfying our assumptions under which price increases sufficiently steeply in λ that physical units sold are declining in λ. But the greater per-unit revenues more than offset the reduced number of physical units sold.

[47] An important caveat is that our model allows for only one dimension of heterogeneity across plants. While this has the virtue of simplicity and tractability, it is also restrictive. If we were to allow for more than one dimension of heterogeneity, as for instance in Hallak and Sivadasan (2009), then the above predictions for the price–plant size elasticities would not hold in general. However, it is possible to derive plausible conditions under which the predictions would continue to hold. In the first variant, if we assume that one heterogeneous parameter, call it φ, reduces unit input requirements (as in equation (6)) and another, call it τ, affects quality conditional on inputs (as in equation (8)), then the input price–plant size elasticity will be positive if the elasticity of φ with respect to τ is not too negative——in particular, greater than $-\frac{b}{2a}$. For the output price–plant size elasticity to be positive, we need an additional condition that the same elasticity is less than $\frac{b}{2a}$; the upper bound is required because if λ increases too quickly relative to τ, then marginal costs of production (and hence prices) will decrease more quickly than quality rises. In the second variant, if we add an additional heterogeneity parameter, υ, multiplying f_{q} in equation (10) then, assuming $\alpha > \frac{1}{\sigma-1}$, a sufficient condition for both the input price–plant size and output price–plant size elasticities to be positive is that the elasticity of υ with respect to λ is not too negative—in this case, greater than $-2a[\alpha(\sigma-1)-1]$. Under these conditions, we would expect the price–plant size elasticities to vary with differences in the scope for quality differentiation (b or α) in ways similar to the model in the main text.

prices, input prices, and plant size, consistent with the empirical patterns documented in Section 3 above. The model as presented above contains just one sector, but it would be straightforward to extend the model to many sectors. If sectors differ in the parameters capturing the scope for quality differentiation (b in the first variant, α in the second) but are otherwise similar, then it is clear from equations (12a) and (12b) and equations (13a) and (13b) that the output price–plant size and input price–plant size elasticities are predicted to be greater in sectors with more scope for quality differentiation.

Finally, in both cases, since the cut-off for entry into the export market is to the right of the cut-off for entry into the domestic market, export status is positively correlated with λ, and we have the same predictions for the correlations of output and input prices with export status as for the correlations with plant size.[48]

Despite the many similarities, there are also important differences between the two variants of the model. Most obviously, the first assumes away spending on fixed quality costs, while the second predicts that more-capable, larger plants make larger fixed quality investments. Another difference is that only in the second variant does the optimal quality choice of plants depend on the scale of the market(s) to which a plant sells (captured in equations (11a–11e) by the Φ term). In principle, these differences should allow one to discriminate between the two variants empirically. The existing literature provides mixed evidence. On the one hand, consistent with the second variant, Bustos (2011) and Lileeva and Trefler (2010) find that exogenous increases in market access to countries of roughly similar income levels lead to greater investments in technology and innovation; although these papers do not focus on product quality, the technology investments may correspond to fixed quality investments as we have modelled them here.[49] On the other hand, a recent paper by Brambilla, Lederman and Porto (2010) compares the effects of exogenous increases of exports by Argentinian firms to Brazil and to richer countries (*i.e.* the U.S. and the European Union) and finds effects on wages only for the latter, consistent with the first variant of our model.[50] In our context, we do not feel that we have either plausible measures of plant-level fixed costs or a convincing source of exogenous variation in market access and hence we focus on the predictions that are common to the two variants and leave for future work the task of testing between them.

5. CROSS-SECTOR EVIDENCE FOR QUALITY INTERPRETATION

In this section, we return to the data and examine the prediction, common to the two variants of our model, that the output and input price–plant size elasticities are increasing in the scope for quality differentiation. The key empirical challenge is to find a measure of the scope for quality differentiation across industries. Here we follow Sutton (1998), arguably the leading existing work on the subject, which uses the ratio of industry-level R&D and advertising expenditures to sales as a proxy for this parameter. This choice can be justified explicitly by the second variant of

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[48] Note that the symmetry between countries in this model implies that if plants enter the export market they will sell the same amount in the export market as in the domestic market, and thus the model does not predict a positive correlation of plant capability and the export share of sales, conditional on exporting. But it is not difficult to imagine extensions to our model in which the export share and plant capability would be positively related, for instance, if capability reduced export costs as well as unit input requirements.

[49] There is also extensive evidence of a positive correlation between exporting and investments in technology and/or R&D in non-quasi-experimental settings; see *e.g.* Aw and Batra (1998), Aw, Roberts and Winston (2007). In other related work, Aw, Roberts and Xu (2011) develop a dynamic structural model with fixed costs of R&D and their estimates suggest that reduction of trade costs (even between similar countries) will lead to an increase in R&D investment as well as exporting.

[50] To be precise, the Brambilla, Lederman and Porto (2010) results are consistent with the first variant of our model extended to allow for greater willingness to pay for quality in richer countries, as in Verhoogen (2008).

our model. From equations (11b) and (11e), note that fixed quality outlays as a share of revenues are constant across plants within an industry:

$$\frac{f_{\mathrm{q}}^{*}(\lambda)}{r^{*}(\lambda)} = \frac{\alpha}{2}\left(\frac{\sigma - 1}{\sigma}\right). \tag{14}$$

Because this ratio does not depend on λ, we know that it also reflects the aggregate value of fixed quality outlays as a proportion of revenues, regardless of the entry cut-offs, λ^* and λ_x^*. Following Sutton (1998), we consider R&D and advertising expenditures to be measures of fixed quality investments. Then assuming a constant elasticity of substitution σ across industries, there is a one-to-one mapping between R&D and advertising expenditures as a share of revenues and the scope-for-quality-differentiation parameter, α.[51] Although the first variant of our model does not posit a role for fixed quality outlays, it does not appear implausible to interpret R&D and advertising intensity as a proxy for the parameter b, on the argument that firms will only invest in R&D and advertising in sectors in which it is possible to affect quality (or perceived quality), and that more capable entrepreneurs are likely to have a greater comparative advantage in using higher-quality inputs in such sectors.

The Colombian manufacturing census does not contain information on R&D or advertising expenditures. Instead we use sector-level information on R&D and advertising intensity from the U.S. Federal Trade Commission (FTC) Line of Business Survey, a source that has been widely used by researchers, including Cohen and Klepper (1992), Brainard (1997), and Antràs (2003), in addition to Sutton (1998). Appendix A2 contains details of the data processing.

A potential concern with using R&D and advertising intensity is that it may reflect *horizontal* differentiation as well as scope for vertical differentiation. Theoretically, it is possible that sectors with greater horizontal differentiation have greater price–plant size elasticities, even holding constant the scope for quality differentiation.[52] To address this concern, we control for horizontal differentiation using two measures. The first is a measure based on Gollop and Monahan (1991) that uses the dissimilarity of *input mixes* across plants within an industry to proxy for horizontal differentiation of outputs.[53] The second measure of horizontal differentiation is the well-known Rauch (1999) measure, based on whether a good is traded on a commodity exchange or has a quoted price in industry trade publications.[54] See Appendix A1 (Supplementary Material) for details on the construction of both measures.

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[51] In the context of a model with a different market structure, Sutton (1998) also demonstrates rigorously that there is a mapping between the (unobserved) scope for quality differentiation in an industry and the (observed) R&D and advertising intensity. See theorem 3.3 and the remark immediately following in Sutton (1998, Chapter 3) .

[52] This pattern holds for instance in the first variant of our model: from (12a) and (12b) it is evident that (as long as $b > 2a$) greater substitutability of goods (a higher σ and hence a higher η) is associated with lower price–plant size elasticities, and conversely that more horizontal differentiation is associated with higher elasticities. Intuitively, in this first variant, more horizontal differentiation does not affect a plant's choice of input quality (see equation (9a)) but it leads to a higher mark-up over marginal costs, which (if $b > 2a$) raises the output price–plant size elasticity. The higher mark-up reduces the elasticity of revenues with respect to capability, λ, and hence raises the elasticity of input prices with respect to revenues. It is worth noting that the same pattern is not predicted by the second variant of our model. While the contrasting predictions of the two variants could in principle be used to discriminate between them—and below the results tend to favour the prediction of the first variant—the contrast reflects a number of special functional form assumptions and we do not interpret the results as offering strong evidence for or against either version.

[53] The Gollop–Monahan index was originally designed to measure diversification across establishments of multi-establishment firms, but it has also been used to measure horizontal differentiation across plants (Syverson, 2004). We are grateful to Chad Syverson for suggesting this measure.

[54] Although it is standard in the trade literature to use the Rauch (1999) index as a measure of horizontal differentiation, arguably it is better interpreted as reflecting differentiation more generally (*i.e.* both horizontal and vertical). Regardless of the interpretation, the key point, documented below, is that the coefficient on the R&D and advertising intensity measure is robust to the inclusion of the Rauch index.

Appendix Table A3 (Supplementary Material) reports summary statistics on R&D and advertising intensity, the modified Gollop–Monahan index, and the Rauch (1999) index, by four-digit industry.[55] The R&D and advertising intensity measure is consistent with a rough intuitive ranking of sectors by scope for vertical differentiation: the most intensive sectors are drugs and medicines, cosmetics and cleaning products, spirits and liquors, office machinery, and tobacco products; the least intensive are cement, tanneries, sugar refining, sawmills, and petroleum refineries.

Before turning to the regression results, we illustrate the key empirical regularities with simple figures. For both outputs and inputs, we estimate a version of our baseline regression equation (1) in which the coefficient on plant size, γ, is allowed to vary flexibly by four-digit sector.[56] We then plot the sector-specific estimates of γ against log R&D and advertising intensity in Figure 3, for outputs in the top panel and inputs in the bottom panel. Each point on the scatterplot indicates a four-digit industry, with weights (total sector output) indicated by the sizes of the plotting symbols. (The same plots with symbols indicating the corresponding sector appear in Appendix Figures A1 and A2.) (Supplementary Material) An obvious feature of both plots is the tremendous variance in estimated price–plant size elasticities that is *not* explained by the R&D and advertising variable. At the same time, a positive relationship between the R&D and advertising variable and the estimated elasticities is discernible for both outputs and inputs.[57]

In our regression specifications, we test whether the price–plant size elasticities are greater in sectors with greater scope for quality differentiation by interacting R&D and advertising intensity with plant size in an equation otherwise similar to equation (1). Based on Figure 3, we expect this interaction term to have a positive coefficient for both outputs and inputs. To control for the effect of horizontal differentiation across sectors, we also include an interaction of one of our measures of horizontal differentiation and plant size.

Table 3 reports the results of regressions including these interactions. For comparison purposes, Columns 1 and 5 report specifications similar to Column 2 of Table 1 for the slightly reduced sample for which all three differentiation measures can be constructed; the point estimates are not statistically different from those in the earlier table. In Columns 2 and 6, we include the interaction of log employment with the R&D and advertising intensity in the output industry of each Colombian plant.[58] To facilitate interpretation, we have deviated the log R&D and advertising intensity variable and the horizontal differentiation measures from their global means prior to interacting them with log employment; the coefficient on log employment

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[55] Not all industries that appear in Appendix Table A2 (Supplementary Material) appear in Appendix Table A3 (Supplementary Material) because of slippage in the concordance process. In Appendix Table A3, we have only included sectors for which both measures of horizontal differentiation are available.

[56] In practice, we interact plant size with a full set of four-digit industry indicators. This method is preferred to estimating equation (1) separately by sector because it takes into account all sales or purchases of a given product in a given year (*i.e.* not just those by plants in a particular four-digit sector) in identifying the product-year effects, ψ_{it}.

[57] An important caveat is that the significant positive slopes in Figure 3 are not robust to all possible weighting schemes. In particular, although the slopes are robust to using total employment, total number of plant–product observations or total number of distinct plants as weights, they are not robust to dropping the weights altogether (*i.e.* giving each sector a weight of one), as in that case several small outlier sectors, which have negligible effect in the weighted regressions, become influential. Caution is thus warranted in interpreting these figures and, more generally, the results of this section.

[58] The definition of output industry differs slightly between Columns 2–4 and 6–8. When the output price is the dependent variable, we define output industry to be simply the first four digits from the eight-digit product code for each plant-product-year observation. When the input price is the dependent variable, this is not possible because particular inputs are not associated with particular outputs. Instead, we use the ISIC four-digit category of the corresponding plant, which is calculated as the industry in which the plant derives the largest share of its revenues. For details, see Appendix A2.

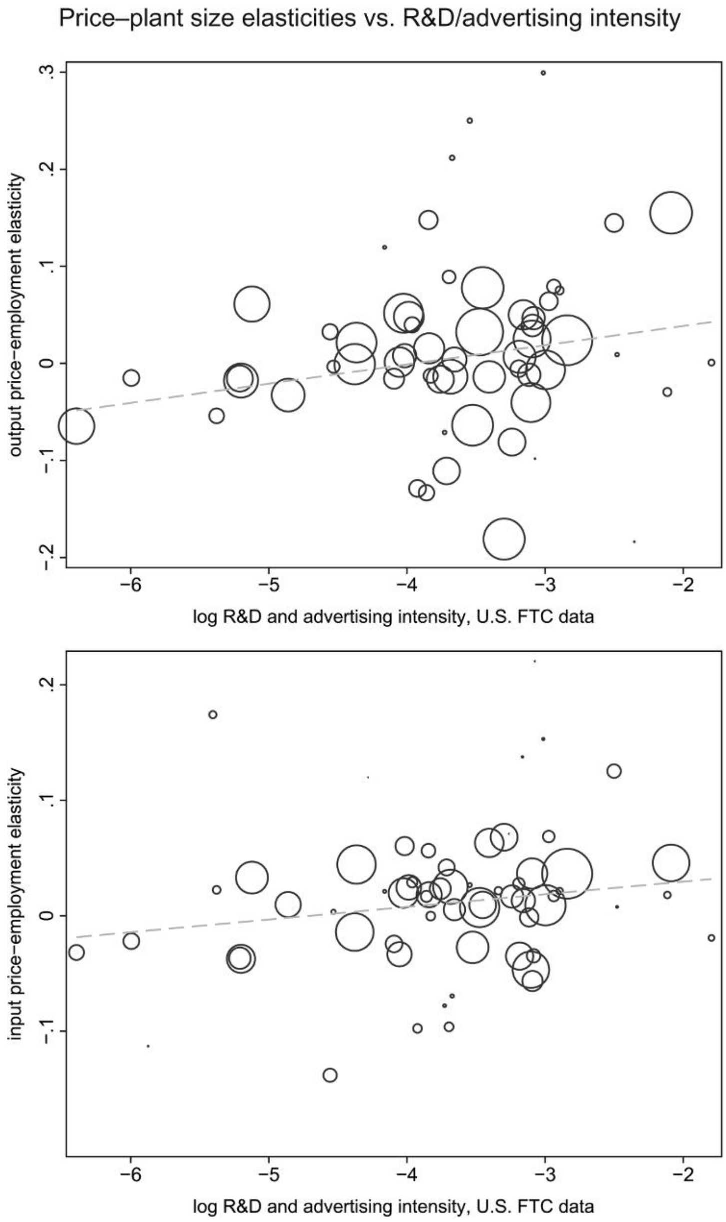

FIGURE 3: Price-plant size elasticities vs. R&D/advertising intensity.

Sector-specific output (input) price–employment elasticities calculated from a regression of log real output (input) unit value on log employment interacted with four-digit sector dummies, with product-year, region-year and five-digit sector dummies as additional covariates. Fitted regression line is weighted by total output in each industry. Weights are indicated by sizes of circles. To increase legibility, the following outliers (with elasticity estimates in parentheses) have been omitted from the graph: 3833 (−0·41), 3530 (−0·38), 3821 (0·72), 3841 (0·83), 3842 (2·46) for outputs, and 3902 (−0·70), 3852 (−0·33), 3530 (0·17), 3853 (0·22) for inputs. These sectors (which have few observations and negligible effects on regression results) are included in the regressions reported in Table 3. To map plotted circles to industries, see Appendix Figures A1 and A2

TABLE 3

Interactions with differentiation measures for output sectors

	Dependent variable: log real output price				Dependent variable: log real input price			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Log employment	0·029*** (0·007)	0·029*** (0·007)	0·028*** (0·007)	0·028*** (0·007)	0·013*** (0·003)	0·013*** (0·003)	0·013*** (0·003)	0·013*** (0·003)
Log emp.*log(R&D + adv. intensity)		0·022** (0·009)	0·022** (0·009)	0·019** (0·009)		0·010** (0·005)	0·009* (0·005)	0·010** (0·005)
Log emp.*Gollop-Monahan index			0·126*** (0·035)				0·051** (0·024)	
Log emp.*Rauch index				0·045*** (0·015)				−0·004 (0·009)
Product-year effects	Y	Y	Y	Y	Y	Y	Y	Y
Industry effects	Y	Y	Y	Y	Y	Y	Y	Y
Region-year effects	Y	Y	Y	Y	Y	Y	Y	Y
R^2	0·90	0·90	0·90	0·90	0·79	0·79	0·79	0·79
N (obs.)	341,812	341,812	341,812	341,812	1,133,366	1,133,366	1,133,366	1,133,366
N (plants)	12,356	12,356	12,356	12,356	11,406	11,406	11,406	11,406

Notes: Data on advertising and R&D expenditures as a share of total industry sales are from the U.S. FTC 1975 Line of Business Survey, converted from FTC four-digit industry classification to ISIC four-digit rev. 2 classification using verbal industry descriptions. At SITC four-digit level, Rauch (1999) measure is set to 0 if good is “homogeneous” or “reference-priced” according to the Rauch “liberal” definition, to 1 if reported not to be in either category, and then concorded to ISIC rev. 2 four-digit categories. Differentiation measures have been deviated from global means. The table uses the 1982–2005 panel. Sample includes sectors with complete data on advertising and R&D intensity, Gollop–Monahan index and Rauch measures, which excludes a small number of sectors. Columns 1 and 6 correspond to Column 2 of Table 1 but use reduced sample. Errors clustered at plant level. N (plants) reports number of clusters (*i.e.* distinct plants that appear in any year). Four-digit industry for outputs (Columns 1–4) defined as first four digits from eight-digit product code. Four-digit industry for inputs (Columns 5–8) defined as industry of corresponding plant. Number of plants differ between Columns 1–4 and 5–8 because of difference in industry definitions. Robust standard errors are given in parentheses. *10% level, **5% level, ***1% level. See Appendix A1 for more detailed variable descriptions and Appendix A2 for details of data processing.

(uninteracted) thus reflects the output price–plant size elasticity in a sector with average values of the differentiation measures. Consistent with the pattern illustrated in Figure 3, both the output price–plant size elasticity and the input price–plant size elasticity are significantly greater in sectors with more scope for quality differentiation.

Columns 3 and 7 of Table 3 control for differences in horizontal differentiation by including the interaction of log employment with the modified Gollop–Monahan index.[59] The estimates of the coefficient on this interaction suggest that the price–plant size elasticities are greater in more horizontally differentiated sectors, consistent with the first variant of our model (see footnote 52.) The more important point, however, is that including the interaction with the Gollop–Monahan measure has little effect on the coefficient estimates for the interactions with R&D and advertising intensity. Columns 4 and 8 report regressions using the Rauch (1999) index. The coefficient on the Rauch term in Column 8 is insignificant, but overall the results are similar to those using the Gollop–Monahan index.

Although not explicitly in our model, it would be plausible to suppose that there are differences in the extent of quality differentiation across intermediate-input sectors as well as across output sectors. In this context, one would expect the input price–plant size correlation to be more positive for inputs purchased from intermediate-input sectors with greater scope for quality differentiation. Table 4 shows that this is the case, using log R&D and advertising intensity and the differentiation measures defined for input sectors, rather than output sectors as above. In particular, the coefficient on the interaction of log employment and log R&D and advertising intensity for the relevant input sector is positive, significant, and robust to the inclusion of controls for horizontal differentiation.

To sum up, we have found that output price–plant size and input price–plant size elasticities are greater in sectors with greater scope for quality differentiation as measured by R&D and advertising intensity in the U.S., consistent with the predictions of our theoretical model. It is worth noting that our estimates are consistent with the existing empirical literature using product-level unit-value data for homogeneous sectors in the U.S., which have reported negative output price–plant size correlations (Roberts and Supina, 1996, 2000; Syverson, 2007; Foster, Haltiwanger and Syverson, 2008). We see in Figure 3 that a number of sectors have negative estimated output price elasticities. Also, the regression results in Table 3 predict a negative correlation for the most homogeneous sectors.[60] At the same time, the results for the average price–plant size correlations in Section 3 suggest that the most homogeneous sectors are not representative of manufacturing overall in Colombia.

6. ALTERNATIVE EXPLANATIONS

Our theoretical model is not the only possible explanation for the empirical facts of Section 3. In this section, we discuss and examine the evidence for what we see as the leading category of alternative explanations: models with plant-specific demand shocks and imperfect competition in input markets.

There are many factors that may lead to greater demand for the products of a particular plant that do not correspond to attributes chosen by the plant that are valued by consumers and hence

[59] Note that any differences in horizontal differentiation that affect prices of all plants equally are already captured by the product-year and industry effects; the key question is whether horizontal differentiation affects the price–plant size elasticities, and that is what the interaction term picks up.

[60] Consider a sector one standard deviation below the mean on both the R&D and advertising intensity and horizontal differentiation measures. The standard deviation of log R&D and advertising intensity is 0·676 and the standard deviation of the Gollop–Monahan index is 0·139. The estimates in Column 3 of Table 3 indicate that the estimated output price–plant size elasticity for the sector will be $-0{\cdot}003$.

TABLE 4
Interactions with differentiation measures for input sectors

	Dependent variable: log real input price			
	(1)	(2)	(3)	(4)
Log employment	0·013*** (0·003)	0·013*** (0·003)	0·013*** (0·003)	0·013*** (0·003)
Log emp.*log(R&D + advertising intensity, input sector)		0·008*** (0·003)	0·008*** (0·003)	0·005* (0·003)
Log emp.*GM index (input sector)			0·008 (0·011)	
Log emp.*Rauch index (input sector)				0·026*** (0·008)
Product-year effects	Y	Y	Y	Y
Industry effects	Y	Y	Y	Y
Region-year effects	Y	Y	Y	Y
R^2	0·76	0·76	0·76	0·76
N (obs.)	1,037,055	1,037,055	1,037,055	1,037,055
N (plants)	13,128	13,128	13,128	13,128

Notes: Differentiation measures are assigned to input sectors, defined as first four digits from eight-digit product code for corresponding input. Sample includes sectors with complete data on advertising and R&D intensity and Rauch measures. Data on advertising and R&D expenditures as a share of total industry sales are from the U.S. FTC 1975 Line of Business Survey, converted from FTC four-digit industry classification to ISIC four-digit rev. 2 classification using verbal industry descriptions. At SITC four-digit level, Rauch (1999) measure set to 0 if good is "homogeneous" or "reference-priced" according to the Rauch "liberal" definition, to 1 if reported not to be in either category, and then concorded to ISIC rev. 2 four-digit categories. Differentiation measures have been deviated from global means. The table uses the 1982–2005 panel. Sample includes sectors with complete data on advertising and R&D intensity, Gollop–Monahan (GM) index and Rauch measures for input sectors, which excludes a small number of sectors. Column 1 corresponds to Column 2, Panel B of Table 1 but uses reduced sample. Errors clustered at plant level. *N* (plants) reports number of cluster (*i.e.* distinct plants that appear in any year). Robust standard errors are given in parentheses. *10% level, **5% level, ***1% level. See Appendix A1 for more detailed variable descriptions and Appendix A2 for details of data processing.

do not correspond to the interpretation of product quality in our model. One example might be favours from a well-placed government procurement official. Another might be collusive agreements between firms not to compete head-on in particular markets. In Dixit-Stiglitz-type demand systems, output prices are fixed multiplicative mark-ups over marginal costs, and since such demand shocks do not affect costs they would not affect output prices. But in frameworks with endogenous mark-ups, one would expect such demand shocks to have effects on output prices.

One such framework, useful for organizing the discussion, is the model of Foster, Haltiwanger and Syverson (2008). The model considers monopolistically competitive firms producing horizontally but not vertically differentiated products, using a demand system similar to Melitz and Ottaviano (2008). In this setting, plant-specific demand shocks unrelated to quality reduce the demand elasticity facing the particular plant and induce the plant both to raise prices and to increase output. This may generate a positive correlation between plant size and output prices even in the absence of quality differentiation.[61]

Regarding the input price–plant size elasticity, Foster *et al.* assume that heterogeneity in input prices is explained by input cost shocks that affect costs without shifting demand and lead

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[61] In this model, there are two offsetting effects. On the one hand, demand shocks tend to induce a positive output price–plant size correlation. On the other, greater productivity tends to induce a negative price–plant size correlation, as in the standard interpretation of the Melitz (2003) model. Which effect will dominate is not clear *a priori*. Empirically, the authors find a negative, insignificant correlation between price and several measures of total output in the homogeneous sectors they focus on (Foster, Haltiwanger and Syverson 2008, Table 1).

to unambiguously lower sales for plants hit by increases in input prices. The model thus predicts a negative relationship between input prices and plant size. But two extensions of the model with imperfect competition in input markets may be consistent with a positive input price–plant size elasticity.

First, plants facing positive demand shocks for their output may face lower price elasticity of output demand, which may in turn lead them to be less sensitive to the prices of inputs. If suppliers have market power, they may optimally charge higher prices to these less price-sensitive producers. Halpern and Koren (2007) present a model with this feature, which they call "pricing-to-firm."[62] This mechanism could predict a positive correlation of input prices and plant size, accompanying the positive correlation of output prices and plant size generated by demand shocks in the output market. Note, however, that it only does so for input sectors in which suppliers have market power.

Second, if final-good producers have monopsony power in input markets and face upward-sloping supply curves for inputs (see *e.g.* Manning, 2003), then a plant-specific demand shock will generate an increase in derived demand for inputs, which will in turn lead plants to move up the input supply curve and pay a higher input price. This mechanism predicts a greater input price–plant size elasticity in input sectors in which buyers are more concentrated or among buyers that have larger "purchaser shares", *i.e.* that purchase a larger share of total domestic production of the input. Note, however, that on the basis of this mechanism we would not expect a positive input price–plant size correlation among purchasers that have no monopsony power and presumably face flat, or very nearly flat, input supply curves.

We now turn to an empirical examination of these alternative hypotheses. Given the richness of the data, we are able to construct measures of market power in intermediate-input markets at a more disaggregated level than is typically possible. We construct two measures of the market power of input suppliers: (1) a Herfindahl index for suppliers of each eight-digit input, defined as the sum of squared market shares of producers of the input; and (2) a modified Gollop–Monahan index defined for the relevant intermediate-input sector (see Appendix A1, Supplementary Material for details). We interpret the latter as a measure of horizontal differentiation of intermediate-input suppliers, which is likely in turn to affect their market power in selling to downstream producers. We construct two measures for the monopsony power of downstream producers in input markets: (1) a Herfindahl index for *purchasers* of each eight-digit input, defined as the sum of squared shares of expenditures on the input by downstream producers; and (2) an individual plant's share of total domestic expenditures on a given input, which we refer to as "purchaser share".[63] The latter measure varies within-product category, unlike the Herfindahl indices; it is likely to be correlated with plant size but not perfectly so since input mixes vary across plants. Appendix Table A4 (Supplementary Material) reports values of the Herfindahl indices averaged across product categories to the four-digit industry level. Because some goods are sold directly to consumers and not purchased by other manufacturing plants, there are a few product categories for which we cannot calculate the Herfindahl purchaser index; these are dropped from the analysis in this section.

Table 5 reports regressions of log input prices on log employment and interactions of log employment with our two measures of supplier market power in input markets.[64] For comparison

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[62] A related idea is that suppliers may have bargaining power vis-a-vis downstream producers, and be able to bargain for a share of the gains from trade, as for instance in Antràs (2003). Bargaining by unions in a labour-market context is analogous.

[63] We are grateful to Andrew Foster for suggesting this measure.

[64] In this table, unlike in Tables 3 and 4, we do not deviate the market-power terms from global means; the coefficients on the uninteracted log employment term thus represent the input price–plant size elasticities in the sectors with the least market power.

TABLE 5

Interactions with market power measures for input sectors

	Dependent variable: log real input unit value					
	(1)	(2)	(3)	(4)	(5)	(6)
Log employment	0·012***	0·019***	0·010*	0·010***	0·009***	0·018***
	(0·003)	(0·004)	(0·006)	(0·003)	(0·003)	(0·007)
Log emp.*Herfindahl index (suppliers)		−0·015**				−0·019***
		(0·006)				(0·006)
Log emp.*GM index (input sector)			0·004			−0·000
			(0·011)			(0·011)
Log emp.*Herfindahl index (purchasers)				0·016		−0·003
				(0·011)		(0·011)
Purchaser share					0·240***	0·249***
					(0·037)	(0·038)
Product-year effects	Y	Y	Y	Y	Y	Y
Industry effects	Y	Y	Y	Y	Y	Y
Region-year effects	Y	Y	Y	Y	Y	Y
R^2	0·76	0·76	0·76	0·76	0·76	0·76
N (obs.)	1,050,029	1,050,029	1,050,029	1,050,029	1,050,029	1,050,029
N (plants)	13,285	13,285	13,285	13,285	13,285	13,285

Notes: Herfindahl index of suppliers is sum of squared market shares of producers of input, at eight-digit industry level. Herfindahl index of purchasers is sum of squared expenditure shares of purchasers of input, at eight-digit industry level. The Gollop–Monahan (GM) index (input sector) is assigned based on industry of inputs. The purchaser share defined as expenditures on product by plant as a share of total expenditures on product by all plants in a given year. Differentiation and concentration measures have not been deviated from global means (unlike in Tables 3 and 4). The table uses the 1982–2005 panel. Sample includes plant-product-year observations for which both Herfindahl indices, GM index and purchaser share could be constructed; plants that use only non-manufacturing inputs are excluded. Errors clustered at plant level. N (plants) reports number of clusters (*i.e.* distinct plants that appear in any year). Robust standard errors are given in parentheses. *10% level, **5% level, ***1% level. See Appendix A1 for more detailed variable descriptions and Appendix A2 for details of data processing.

purposes, Column 1 reports the results from our baseline regression on the slightly reduced sample; the estimated input price–plant size slope is similar to the estimates above.

In the pricing-to-firm story discussed above, one would expect to see a greater input price–plant size elasticity in input sectors in which suppliers have more market power. In Column 2, the coefficient on the interaction of the Herfindahl supplier index and log employment is *negative* and significant. That is, the input price–plant size elasticity is *lower* in input sectors in which suppliers are more concentrated. One possible explanation for this result is that larger plants may have more bargaining power with suppliers than smaller plants, and that this matters particularly in sectors in which suppliers have market power; an analogous explanation that has been advanced about the role of Wal-Mart in the retail sector of the U.S. (see *e.g.* Basker, 2007). In Column 3, we include an interaction of the Gollop–Monahan index defined for the relevant intermediate-input sector, and the coefficient on the interaction is statistically indistinguishable from zero. In short, there is little evidence that the input price–plant size correlation is more positive in input sectors in which input suppliers have more market power, which argues against the pricing-to-firm story in this context.

In the monopsony story discussed above, downstream producers facing upward-sloping input supply curves would pay higher input prices as a consequence of positive demand shocks for their output. There is some evidence to support this story. The coefficient on the interaction of log employment with the Herfindahl *purchaser* index is positive (but not significant) in

Column 4, and the coefficient on the interaction with the purchaser share in Column 5 is positive and significant. That is, plants with a larger share of total purchases of a particular input pay a higher price for the input. Column 6 shows that this result is robust to including all concentration measures simultaneously.

However, the key point of the table is that the coefficients on the *uninteracted* log employment term remain positive and significant across specifications. These results can be interpreted as indicating that in input sectors where suppliers and/or purchasers have approximately zero market power, the positive input price–plant size correlation remains. It is also worth noting that explanations based on market power would not predict the systematically greater price–plant size elasticities in sectors with more scope for quality differentiation that we observed in Section 5.[65] Overall, we interpret the results as suggesting that observable measures of market power in input markets do not provide a complete explanation for the positive input price–plant size correlation.

7. CONCLUSION

This paper has used uncommonly rich and representative data on Colombian manufacturing plants to document several new facts: on average, within narrowly defined sectors, there are positive correlations between output prices and plant size and, perhaps more surprisingly, between input prices and plant size. Similar patterns hold for the relationships between prices and export status. These correlations vary systematically across sectors: using an off-the-shelf measure of the scope for quality differentiation from Sutton (1998)—R&D and advertising intensity of industries in the U.S.—we have found that the output price–plant size and input price–plant size elasticities are greater in sectors with more scope for quality differentiation.

These empirical patterns are consistent with a parsimonious extension of the Melitz (2003) model to allow for the endogenous choice of input and output quality. We have developed two variants of the model, which differ in the production function for quality. In the first, plant capability and input quality are complements in generating output quality, and there are no fixed costs of improving quality. In the second, there are fixed costs of improving quality and producing high-quality outputs also requires high-quality inputs, but no direct complementarity between plant capability and input quality. The two variants have similar observable implications, and in particular predict (conditional on positive scope for quality differentiation) a matching between more capable entrepreneurs and higher-quality inputs in producing higher-quality outputs in equilibrium.

We have considered alternative explanations based on differences in market power. While we have found some support for the hypothesis that plants have monopsony power in input markets, it does not appear that that market-power differences can account fully for the empirical patterns. Overall, although we do not observe product quality directly and the evidence is therefore not definitive, we interpret the empirical results as suggestive of an important role for quality differences of both inputs and outputs in explaining the plant-product-level price dispersion.

In addition to the implications mentioned in the introduction, this conclusion carries potentially important implications for the literature on the productivity effects of international integration. In this literature, it has been standard to use TFP as the primary measure of plant or firm performance. But standard methods for estimating TFP assume that neither outputs nor inputs

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[65] In results available from the authors, we include an R&D and advertising intensity–plant size interaction together with the Herfindahl purchaser index and the purchaser share measure, and find that the coefficients on the R&D and advertising term remain positive and significant.

are differentiated on a quality dimension. If such quality heterogeneity exists, then standard TFP measures are likely to be subject to both upward and downward biases, with ambiguous net effects (Katayama, Lu and Tybout, 2008; De Loecker, forthcoming). A number of existing results on the productivity effects of exporting may need to be re-evaluated in this light.

We conclude with a caveat about external validity: the manufacturing sector of Colombia may not be representative of manufacturing sectors elsewhere. Although there is some evidence that similar patterns hold in other countries at approximately Colombia's level of development,[66] it is not clear to what extent they hold among plants in richer countries, which presumably tend to produce in higher-quality, higher-income segments of the product spectrum. One would not be surprised to learn that the positive price–plant size correlation does not hold for producers of French wines or Swiss watches, for instance.[67] The extent to which the findings for Colombia generalize to other countries remains an open question, to which one hopes the increasing availability of product prices in plant-level data sets will soon provide an answer.

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Supplementary Data

Supplementary data are available at *Review of Economic Studies* online.

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[66] As noted above, an early draft of Hallak and Sivadasan (2009) documented a positive price-plant size relationship in Indian data, and Verhoogen (2008) finds that larger plants in Mexico were more likely to have ISO 9000 certification, an international production standard commonly interpreted as a measure of product quality.

[67] Holmes and Mitchell (2008) make an analogous argument that the plant size–skill correlation may be of different signs in different historical contexts.

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