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Inferring market power under the threat of entry: the case of the Brazilian cement industry

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Consider a setting where threatened rather than actual import competition restrains a domestic oligopoly's prices. I show that not modelling the entry threat may underestimate the true degree of market power, as incumbents' blunted price responses to demand shocks resemble perfectly competitive behavior. Evidence from Brazilian cement markets points to an important role for imports in determining domestic cement prices, despite the near absence of imports. On assuming autarky, models with market power are rejected in favor of competition among incumbents. However, allowing a role for imports rejects the autarky assumption and precludes one from rejecting the presence of market power.

1. Introduction

■ Industrial organization researchers and antitrust practitioners have long been concerned with supply-side estimation with a view to measuring, in the absence of cost data, the degree of market power exercised by firms, or their price-cost margins. A stream in the literature, dating back to Bresnahan (1982) and Porter (1983), estimates a parametric pricing equation that nests a class of standard oligopoly models such as monopoly, Cournot, and competition. As in Porter's work, a key assumption underlying the structural researcher's inference is that one of these theoretically consistent oligopoly models governs the data-generating process.[1] In a more recent stream (Gasmi,

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[1] Given the structure, identification between marginal cost and the markup works through demand, as the response of prices to changes in demand conditions depends on whether firms enjoy market power. See Reiss and Wolak (2005) for a thorough discussion, including interpretation, of this framework.

Laffont, and Vuong, 1990), the behavioral interpretation is explicit. The researcher departs from a menu of (possibly nonnested) oligopoly models that are deemed plausible *a priori* and estimates supply conditioning on each model's solution concept; nonnested tests (e.g., Vuong, 1989; Smith, 1992) are then used to select among the competing models.

This article extends this literature by considering a setting where one such coherent oligopoly theory—monopoly, say—is the true behavioral model generating prices, but prices are censored. In particular, a domestic oligopoly faces the threat of import competition. I model a competitive fringe of elastically supplied imports, where delivered marginal cost—a world price plus the inbound trade cost—is high enough that imports do not occur in equilibrium, but low enough that it limits the domestic oligopoly's prices in some markets (say temporal markets, where the local currency is strong so that imports are at their most competitive).[2] Like its domestic counterpart, the cost of imports is not observed (by the researcher). Despite the binding price ceiling, incumbents' price-cost margins are large. The researcher wishes to estimate these.

I begin by showing that were the researcher to overlook the constraining effect of potential entry on a subset of observations, and thus not account for the censoring condition, he might obtain a downward-biased estimate of the true price-cost margins, inferring that domestic outcomes are "more" competitive when they are less. He might, for example, commit a type I error by rejecting the true hypothesis of (constrained) market power in favor of the alternative of domestic competition and zero price-cost margins. Or the researcher might commit a type II error by failing to reject the false hypothesis of zero price-cost margins. Intuitively, the threat of foreign entry acts to constrain the ability of firms with market power to respond to exogenous demand shocks.

To drive the point home, consider the following polar example. Suppose demand is perfectly inelastic at some quantity q up to a threshold price p^*, above which it is zero. A domestic monopolist's marginal cost is zero. The monopolist will charge the marginal cost of imports $c^I < p^*$ regardless of any demand fluctuations, be they shifts in p^* or q. Whereas this looks exactly like competitive behavior with marginal cost of c^I, a researcher inferring this would be grossly understating the markup, which is in fact infinite.

In the spirit of Lee and Porter (1984), I next show how the researcher can account for the "latent" foreign sector. By modelling the endogenous switching mechanism between unconstrained and constrained markets, supply parameters—domestic and foreign—can be consistently recovered from censored market outcomes (maintaining the structural assumption that the domestic supply component is appropriately specified).

I examine the Brazilian cement industry, investigating leads from industry sources that, minimal import quantities notwithstanding,[3] the threat of imports places downward pressure on domestic cement prices. Several features of the data are consistent with the latent import competition hypothesis. Notably, observed variables that move the latent price of cement imports—primarily the exchange rate (and large swings in its real value have occurred over the sample period)—explain much of the variation in observed domestic cement prices. This occurs over and above observed domestic cost-shifters (as well as observed local demand shocks): the domestic industry's input prices are only imperfectly correlated with the exchange rate or the world oil price. Also, I interpret the low market price elasticities of demand that I systematically (and consistently) estimate across Brazil's local consumer markets—of the order of -0.5, including markets where there is essentially one seller—as further corroborative evidence of a price limit in the industry. Of course, inelastic market demand in equilibrium would also be consistent with competition among incumbents. But the domestic marginal costs that I am able to construct,

[2] I refer to this as the "*latent* imports-as-market-discipline" hypothesis (to borrow from Levinsohn, 1993), by which international trade can (partially) restrain local market power without any trade taking place. (I thank a referee for emphasizing this point.) For instance, in a widely cited study of a U.S. sugar cartel, Genesove and Mullin (1998) document how the CEO of the Sugar Trust (the market leader with a 63% share) "acknowledged setting the price of refined sugar so that none would be imported from Europe."

[3] Imports have accounted for 1%–2% of Brazilian cement consumption. By comparison, the United States has imported as much as 30% of its domestic cement needs (and sourced from around the world).

by virtue of an unusually detailed data set and the simple technology of cement, indicate that price-cost margins are far from competitive, amounting to about 50% of producer prices.

Taking costs as unknown, I then estimate the supply side either under the autarky assumption or allowing for the latent effect of trade integration. First, I assume away a foreign sector, effectively imposing the unconstrained regime across the sample. On specifying a standard (constraint-free) pricing equation with markup parameter (as in, e.g., Porter, 1983), I estimate price-cost margins that hover around zero with tight confidence intervals, suggesting a competitive industry. Similarly, on adopting a menu approach (as in, e.g., Gasmi, Laffont, and Vuong, 1992), Vuong (1989) selection tests indicate that perfect competition strongly outperforms monopoly and Cournot in explaining the data. Were the researcher to depart from this menu of *autarkic* oligopoly models, he would reject market power in favor of zero price-cost margins.

Second, in light of the institutional analysis, I incorporate the import price ceiling in the supply side of the structural model. In addition to market power, the exercise provides an empirical test of the autarky assumption. Estimating supply in the markup parameter tradition (i.e., now embedding, in a mixture model, an imports pricing equation alongside the earlier domestic pricing equation), I am not able to recover significantly higher price-cost margins than before, and domestic supply estimates are less precise. It may well be that the relatively unsophisticated domestic pricing component of the enlarged trade model does not adequately capture any dynamics in the data (e.g., as in Corts, 1999). Where the model shows its potential is on the imports pricing side. In contrast to their domestic counterparts, imports supply parameters are tightly estimated, with the coefficient on the exchange rate taking center stage in terms of economic and statistical significance. Imports matter quantitatively, no matter their minimal quantity. Returning to the menu approach, I expand the earlier set of autarkic formulations to include models of market power that allow for latent import competition, namely domestic monopoly or Cournot facing a price ceiling. I now find that either *integrated* model of market power outperforms (autarkic) perfect competition. (And I am unable to discriminate between integrated domestic monopoly and integrated domestic Cournot: conditional on either solution concept, the data suggest that virtually all markets in the sample are constrained.) The model selection tests lead me to conclude that *by allowing imports to play a role*: (i) the hypothesis that Brazilian cement prices are determined in autarky is rejected in favor of a role for imports, and (ii) the presence of market power can no longer be rejected.

The plan of the article is as follows. Section 2 considers the estimation of supply under the threat of international competition. Section 3 presents the empirical application, by way of institutions, demand, and supply. Section 4 concludes.

2. Inferring markups under the threat of entry

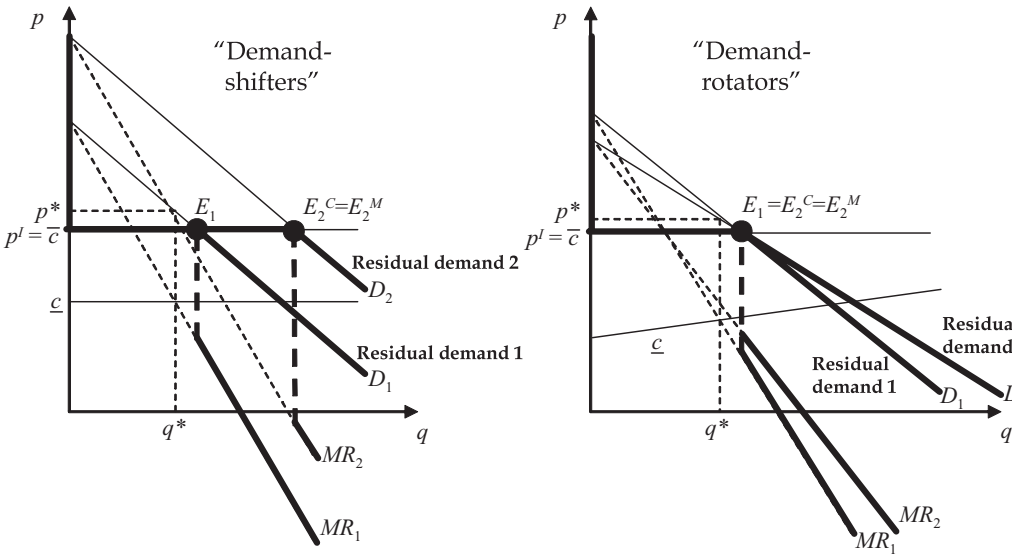
■ I extend the setting considered by Bresnahan (1982) and Porter (1983) to a context where a domestic oligopoly—a local monopoly, say—is constrained by international competition in a subset of markets. The domestic monopolist, with low marginal cost $\underline{c}$, faces a competitive fringe of imports (labeled I) with perfectly elastic supply at landed price (cost) p^I. This delivered price of imports is high enough relative to domestic marginal cost—so in equilibrium there is no foreign entry—yet at the same time low enough to occasionally restrain the incumbent from setting the autarkic monopoly price p^*.[4] The observed price is thus given by $p = \min(p^*, p^I)$, where p^* and p^I are latent.

Figure 1 illustrates how Bresnahan's (1982) identification result no longer holds. In the presence of a binding price ceiling, neither shifts (left panel) nor rotations (right panel) in demand elicit differential price responses that would reveal the degree of market power. From Bresnahan's example, the alternative models of (i) a competitive industry with high cost $\overline{c}$, and (ii) a monopoly

[4] This is plausible in the context of a large world market, where inbound trade costs are added to a world price. Should the fringe enter and be observed, a researcher would naturally incorporate it in the structural model (as, for example, Suslow, 1986 does).

FIGURE 1

PRICE-COST MARGINS ARE NO LONGER IDENTIFIED WHEN THE ENTRY THREAT BINDS



In the left panel, demand shifts (and marginal cost is flat in output). In the right panel, demand rotates (and marginal cost varies in output).

(cartel) with low cost $\underline{c}$ facing a price limit at $\bar{c} \equiv p^I$, are not observationally distinct. Unaware of the latent ceiling, a researcher might inadvertently take the lack of price variation to exogenous movements in demand as evidence in the direction of price-taking behavior and zero price-cost margins. (Notice that around a "constrained equilibrium," demand is still identified: fluctuations in the import price, likely correlated with domestic cost-shifters often used as instruments, move the kinked equilibrium up and down along the market demand curve.)

The empirical literature on market power typically assumes that the observed price p and output q solve the system given by the market demand function

$$q = D(p, Y, \varepsilon^d; \alpha) \tag{1}$$

and the constraint-free (i.e., autarkic) static pricing equation

$$p + \theta q \frac{\partial p(q, Y, \varepsilon^d; \alpha)}{\partial q} - c(q, W; \beta) - \varepsilon^s = 0, \tag{2}$$

where $p(q, .) \equiv D^{-1}(q, .)$ is inverse demand, $c(q, .)$ is marginal cost, Y and W are, respectively, observed exogenous demand and cost covariates, mean-zero error terms ε^d and ε^s, respectively, capture the unobserved exogenous components of demand and supply, and (α, θ, β) are unknown parameters to be estimated.[5] I formalize the inconsistency argument using the nested "markup" (or "conduct") parameter approach of Bresnahan (1982) and Porter (1983), with the θ parameter limited to interpretable values, such as monopoly or perfect collusion ($\theta = 1$), symmetric Cournot (θ being the reciprocal of the number of firms), and perfect competition ($\theta = 0$)[6]; I will illustrate in the subsequent application how the argument carries through to the nonnested model selection approach of Gasmi, Laffont, and Vuong (1990). Rearranging (2), one obtains the elasticity-adjusted price-cost markup:

[5] Examples are too numerous to cite comprehensively. The aggregate equation (2) is an average across firm-level pricing equations; the latter can alternatively be used to specify the supply side.

[6] To emphasize, I strictly interpret (2) as nesting specific theoretically consistent oligopoly models.

$$\theta = -\eta \frac{p - c}{p}, \tag{3}$$

where $\eta(q, .) \equiv (p(q, .)/q)(\partial p(q, .)/\partial q)^{-1}$ is the market price elasticity of demand (and where, for simplicity, arguments have been suppressed and ε^s has been subsumed into c). This price-cost markup is the object that the researcher wishes to estimate. Following (or simultaneous with) the estimation of demand (1)—which yields a consistent estimate of $\partial p(q, .)/\partial q$—supply specification (2) is estimated by instrumental variables (IV) or the general method of moments (GMM). The identifying assumption is $E(Y'\varepsilon^s) = 0$, where excluded demand covariates Y serve as instruments for the endogenous elements of $c(q, .)$ and $q(\partial p(q, .)/\partial q)$.

With international competition, the researcher no longer observes the solution to demand (1) and supply (2): denote this latent solution by (p^*, q^*). Rather, the observed price is limited according to the data-generating process (DGP):

$$\begin{aligned} p &= \min(p^*, p^I) \\ &\equiv \min\left(-\theta q^* \frac{\partial p(q^*, Y, \varepsilon^d; \alpha)}{\partial q} + c(q^*, W; \beta) + \varepsilon^s, c^I(W^I; \beta^I) + \varepsilon^I\right), \end{aligned} \tag{4}$$

where competitive imports are priced at marginal cost c^I (like c, this is unobserved), W^I are observed exogenous imports cost-shifters (likely correlated with domestic cost-shifters W, albeit imperfectly), β^I are unknown parameters, and ε^I is an imports pricing equation error. I assume that $c^I(.) > c(D(c^I(.), .), .)$ for all market realizations, so that import competition is latent (as in Figure 1). A market outcome is thus given by the censored price p and the corresponding quantity on the demand curve $q = D(p, Y, \varepsilon^d; \alpha)$.

The researcher wishes to infer price-cost markups from a sample of $t = 1, \ldots, N$ observations, where constrained markets are mixed with unconstrained ones. I first consider the situation where he is not aware of—or does not account for—the price ceiling that binds on a subset of observations. I then consider how the researcher can estimate the DGP (4), obtaining consistent markups. Let the indicator function $\chi := \mathbf{1}[p^* \leq p^I]$ denote the latent regime, such that $\chi = 1$ labels a market equilibrium which is unconstrained by the threat of entry and $\chi = 0$ is the event that a market is constrained.

□ **Imposing autarky.** Say that the researcher specifies supply as in Bresnahan (1982) or Porter (1983), which are a limiting case to the DGP (4) with $\Pr(\chi_t = 1) \to 1$. In contrast to the more general true model, the estimated model is

$$p = -\theta q \frac{\partial p(q, Y, \varepsilon^d; \alpha)}{\partial q} + c(q, W; \beta) + \xi^s, \tag{5}$$

where the error of the (mis)specified autarkic pricing equation is denoted ξ^s. The estimated model does not nest the supply decisions of an industry with pricing power that faces a price ceiling; the implication is stated in the following proposition.

Proposition 1. Assume that (i) the exogenous demand variables Y are positive, (ii) the marginal cost function $c(q, W; \beta)$ is linear in parameters β, and (iii) any exogenous additive variable of $c(q, W; \beta)$ is positive. It follows that when the threat of entry constrains prices in an industry that exercises market power, the residual ξ^s in the standard autarkic pricing equation (5) is negatively correlated with Y, that is, $E(Y'\xi^s) < 0$. Consequently, IV estimation of the autarkic pricing equation using demand perturbations yields inconsistent estimates of market power and cost.

Proof. See the Appendix.

Given the inconsistent parameter estimates that result from the failure of the orthogonality condition, $E(Y'\xi^s) < 0$, I investigate the sign of the bias in the estimated markup in a series of

Monte Carlo experiments. (The Appendix reports results for one of these simulated industries.[7]) These simulations confirm the intuition provided above (e.g., Figure 1) that suggests a downward bias in the inference of market power. For both constrained and unconstrained markets, I find that the 95% confidence interval (C.I.) for the (elasticity-adjusted) estimated price-cost markup $-\eta(p-\widehat{c})/p$ overwhelmingly lies below the true realized markup $-\eta(p-c)/p$ (which itself, in a constrained market, lies below the latent markup $-\eta^*(p^*-c^*)/p^* \equiv \theta$, from (3)). That is,

$$-\eta(q,.)\frac{p-c(q,W;\widehat{\beta})}{p} < -\eta(q,.)\frac{p-c(q,W;\beta)}{p} \leq -\eta(q^*,.)\frac{p^*-c(q^*,W;\beta)}{p^*},$$

namely, estimated markup "$\widehat{\theta}$" < realized markup "θ^{actual}" $\leq \theta$ ($\equiv 1$, say in monopoly).

□ **Modelling the import threat.** Suspecting that a price ceiling may be censoring the data, the researcher can model the DGP (4), thus (probabilistically) separating the unconstrained outcomes (where $\chi_t = 1$) from the constrained ones (where $\chi_t = 0$). The supply parameters (domestic and foreign) can then be consistently recovered through maximum likelihood (ML). Specifically, the estimation of domestic supply parameters requires a sufficient number of unconstrained markets in the sample, that is, $\Pr(\chi_t = 1) \gg 0$ (see the Appendix).

In what follows, I condition the likelihood of observing the limited dependent variable $p = p_t$ on the (inferred) curvature of demand. A consistent estimate of $\partial p/\partial q$ can be obtained by two-stage least squares (2SLS) in an earlier stage, and then bootstrapped standard errors can be calculated for the supply estimates to account for the sampling variation in demand. (An alternative to this two-step procedure is to estimate demand along with supply by ML.) Make the distributional assumption that the mean-zero supply shocks $\varepsilon^s, \varepsilon^I$ are i.i.d. normal, with respective variance σ_s^2, σ_I^2. Let $\omega := \{\theta, \beta, \sigma_s, \beta^I, \sigma_I\}$ denote the vector of supply parameters to be estimated, and write[8]

$$[\chi = 1, \text{unconstrained}] \qquad \varepsilon^s = p + \theta q\widehat{\partial p/\partial q} - c(q,W;\beta) \equiv S\left(p, q, \widehat{\partial p/\partial q}, W; \theta, \beta\right)$$

$$[\chi = 0, \text{constrained}] \qquad \varepsilon^I = p - c^I(W^I;\beta^I) \equiv I(p, W^I;\beta^I).$$

As the events $\chi = 1 \Leftrightarrow p = p^* \leq p^I$ and $\chi = 0 \Leftrightarrow p = p^I < p^*$ are mutually exclusive, the (conditional) likelihood of observing $p = p_t$ is

$$L\left(\, p = p_t | q_t, \widehat{\partial p/\partial q}_t, W_t, W_t^I; \omega\right) = \Pr(\chi_t = 1 \cap p = p_t) + \Pr(\chi_t = 0 \cap p = p_t),$$

where, denoting the determinants of the appropriate Jacobians by $|J^s|$ and $|J^I|$ (to account for the transformation in random variables[9]),

$$\begin{aligned}\Pr(\chi_t = 1 \cap p = p_t) &= \Pr\left(p^I \geq p_t\right)\Pr(p^* = p_t)\\ &= \Phi\left(\frac{-I\left(p_t, W_t^I;\beta^I\right)}{\sigma_I}\right)\phi\left(\frac{S(p_t, q_t, \widehat{\partial p/\partial q}_t, W_t;\theta,\beta)}{\sigma_s}\right)\frac{1}{\sigma_s}|J^s|\end{aligned}$$

and

$$\begin{aligned}\Pr(\chi_t = 0 \cap p = p_t) &= \Pr(p^* > p_t)\Pr\left(p^I = p_t\right)\\ &= \Phi\left(\frac{-S(p_t, q_t, \widehat{\partial p/\partial q}_t, W_t;\theta,\beta)}{\sigma_s}\right)\phi\left(\frac{I\left(p_t, W_t^I;\beta^I\right)}{\sigma_I}\right)\frac{1}{\sigma_I}|J^I|.\end{aligned}$$

[7] Other simulations—in which I vary demand and supply specifications while satisfying the assumptions in the proposition—are reported in an online appendix. The online appendix also derives, for a restricted (and algebraically simple) case, an analytical expression for the asymptotic IV estimator for θ, plim $\widehat{\theta}$, in the estimated model (5).

[8] Thus, $S(.)$ provides the ε_t^s that explains p_t conditional on $\chi_t = 1$ and the domestic supply parameters (besides data and demand). Similarly, $I(.)$ conditions on $\chi_t = 0$.

[9] In the imports pricing equation, because $c^I(.)$ is exogenous and thus $\partial\varepsilon^I/\partial p^I = 1$, $|J^I| = 1$.

Note that the implied probability that a market t is unconstrained is given by the relative contribution of the first term to the likelihood, that is, $\Pr(\chi_t = 1|.) \equiv \Pr(\chi_t = 1 \cap p = p_t)/L(p = p_t|.)$. The log likelihood function is then

$$\log L = \log\left(\prod_{t=1}^{N} L\left(\, p = p_t|\, q_t, \widehat{\partial p/\partial q}_t, W_t, W_t^I; \omega\right)\right)$$
$$= \sum_{t=1}^{N} \log L\left(p = p_t|\, q_t, \widehat{\partial p/\partial q}_t, W_t, W_t^I; \omega\right). \tag{6}$$

The present context is reminiscent of disequilibrium models à la Goldfeld and Quandt (1975) and Kiefer (1979), where observed quantity is given by min(Demand(p, X^d, ε^d), Supply(p, X^s, ε^s)), and covariates X and prices p are observed whereas ε is unobserved.

3. Empirical application

■ The empirical setting that I now consider is one where both institutional and econometric evidence point to an important role for imports in determining a domestic oligopoly's prices, despite the virtual absence of actual imports.

□ **Industry: imports discipline hypothesis and data.** By the late 1990s, Brazil ranked sixth among cement-producing countries. The implementation of the Real economic stabilization plan in July 1994 had led to strong consumption growth across the economy, thanks to the reduction in inflation tax brought about by the sharp slowdown in inflation.[10] In the construction sector, notably in housing, this income shock had lifted cement consumption from around 25 million tons per annum (mtpa) in the first half of the decade to 40 mtpa by 1998–1999. The domestic cement industry had consolidated from 19 firms in 1991 to 12 firms by 1999, with the largest producer, Votorantim, accounting for 41% of shipments nationwide.

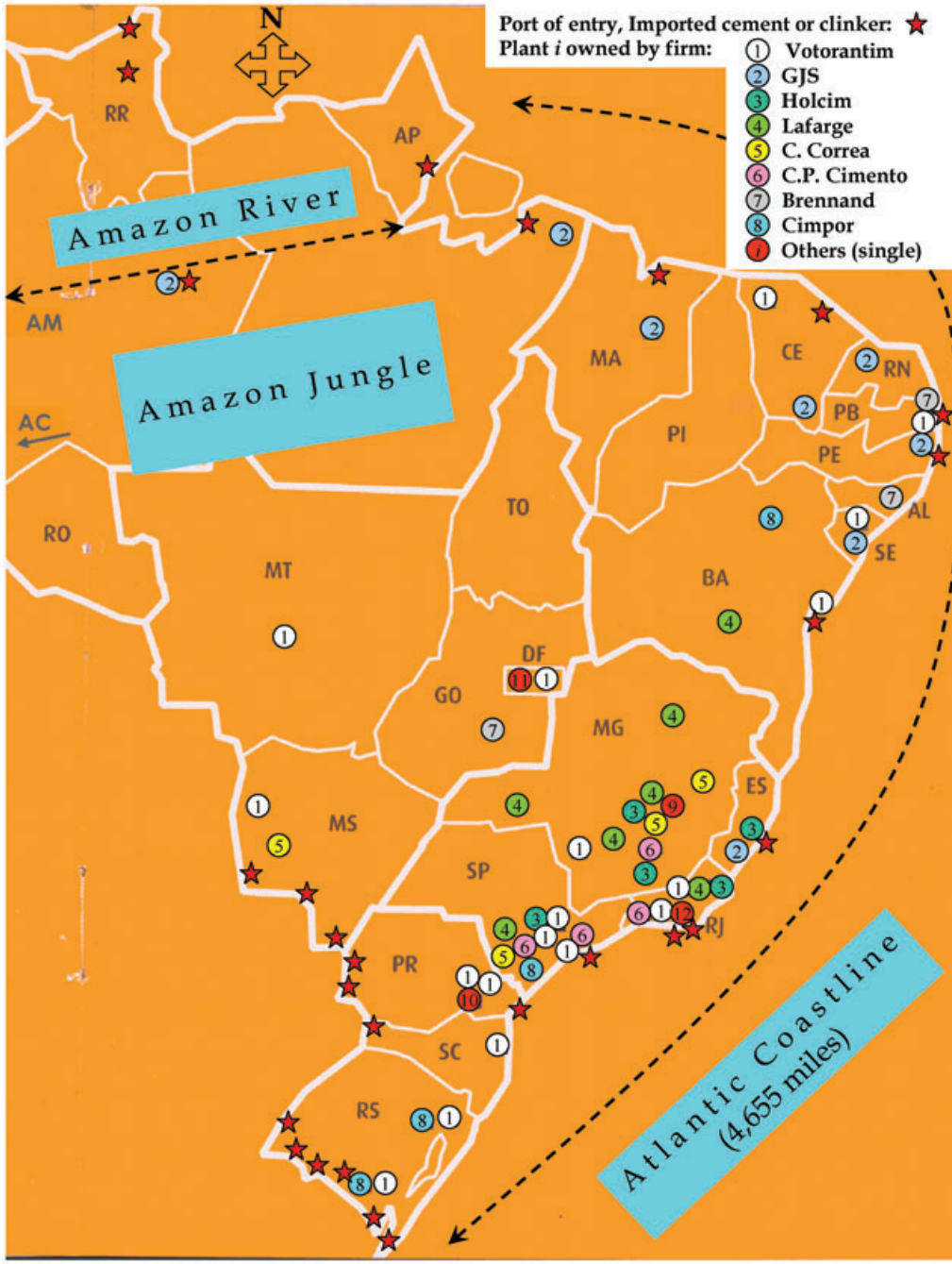
In the wake of Brazil's wide-reaching trade-liberalizing reforms that started in 1990, one might have expected the unpredicted construction boom of 1995–1997 to lead to sizable imports of cement. In the cement industry, new capacity takes years to come on stream, and an international trading market is well established. To illustrate, cement imports into the United States (including a small component in the form of the intermediate product clinker) have accounted for 10%–30% of domestic consumption since the 1980s, with the U.S. Geological Survey (2000) describing "both clinker and cement (as) widely traded."[11] Most of Brazil's population lived in the coastal states, which were the natural candidates for imports to penetrate; the northwestern states were sparsely populated and largely covered with jungle. (This is illustrated in Figure 2 by the spatial distribution of cement plants, which typically locate close to consumers.) Further, the strong appreciation of Brazil's local currency—the Real, denoted R$—between 1994 and 1998 had made imports more competitive; this "exchange rate anchor," a feature of the stabilization plan and which lasted until the devaluation of early 1999, had contributed to a surge in imports across many sectors, from cars to electronics to nondurable consumer goods. However, Brazilian imports of cement and clinker remained subdued, hovering around 1%–2% of domestic consumption during this period and beyond. (Figure 2 also depicts those ports of entry through which cement imports were recorded, even if in small quantity, over an extended period.) The domestic oligopoly was

[10] In particular, the large mass of unbanked low-income households, who had no access to price-indexed checking accounts that protected against inflation, saw a considerable rise in real incomes.

[11] The United States imports cement from around the world (e.g., in 1999, Asia accounted for 38%, the Americas 34%, and Europe 27%). Dedicated seaborne handling and transportation equipment developed in the 1970s gave rise to global trade (Dumez and Jeunemaıtre, 2000). Salvo (2010b) compares the United States to Brazil. For completeness in what follows, Brazil has exported even less cement than it has imported.

FIGURE 2

ACTIVE PLANTS (1999) AND PORTS OF ENTRY (1994–2006)



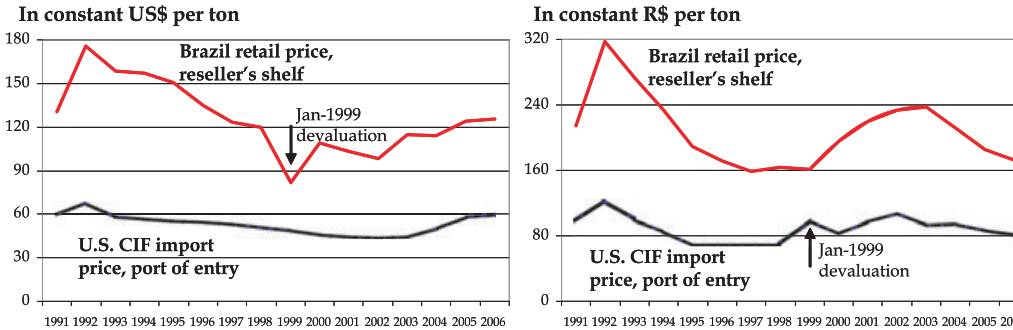
Sources: SNIC, SECEX.

able to meet the exogenous jump in demand, keeping imports at bay, by investing in idle capacity (capacity utilization has historically averaged 65%) and, arguably, by limit pricing.

Paling in comparison to the United States, the limited penetration of imports into Brazil's local cement markets masks a plausible role for imports in restraining domestic prices. This

FIGURE 3

EVOLUTION OF AGGREGATE CEMENT PRICES



Price to the end consumer in (17 "nonjungle" states of) Brazil and CIF import price at (on average 14 East or Gulf Coast) U.S. ports of entry. In the left panel, data series are expressed in constant US$ per ton (U.S. CPI base December 1999). In the right panel, series are in constant R$ per ton (Brazil GPI base December 1999). *Source:* IBGE, US GS.

hypothesis was discussed during several field interviews,[12] and is backed by anecdotal evidence (see Salvo, 2010b for details[13]). Equity analyst Zaghen (1996), for example, wrote: "(a)lthough (cement) imports accounted for only 1.6% of total Brazilian consumption in 1995, reaching 451.3 thousand tons, they represent a constant threat to domestic producers, pressing down domestic prices and imposing a price ceiling of US$70 per ton" (p. 24). The price Zaghen referred to was the cost, insurance, and freight (CIF), at the port of entry. Figure 3 compares the evolution of the aggregate price to end consumers in Brazil against the lower, yet correlated, (annual) CIF import price in the import-heavy United States. (The two series are shown in (i) constant US$ per ton, U.S. consumer price index (CPI) base December 1999; and (ii) constant R$ per ton, Brazil general price index (GPI) base December 1999.) In the late 1990s, Brazil's government even bowed to pressure from domestic producers and passed antidumping measures against specific Venezuelan and Mexican producers, underscoring the threat posed by imports.

Importantly, several features of the disaggregate price data are consistent with the imports discipline story. Table 1 explores how well variables which move the marginal cost of imports—imports cost-shifters W^I—explain variation in domestic cement prices p, comparing their explanatory power with that of domestic cost-shifters W. The reported price regressions (p projected on W^I and/or W) are based on a panel of 17 "nonjungle" states (essentially coastal, as this is where consumers were) over the period January 1991 to December 2006, and all prices are in constant R$, Brazil GPI base December 1999. Specifically, cement prices are shelf prices for the standard 50 kg bag at retailers, known as "resellers."[14] Variables specific to W^I are[15]:

[12] These included representatives of: (i) the domestic cement industry (executives, salespeople); (ii) the industry's trade association (SNIC and its technical arm ABCP); (iii) importers (executives at independent imported clinker grinders Mizú, Davi, and Cimec, as well as an international trader who actually imported cement); (iv) cement buyers (retailers, ready-mix concrete firms, construction firms, concrete product manufacturers); and (v) local construction-sector trade associations (SINDUSCONs).

[13] To describe one instance, for what it is worth, a high-ranking industry officer (whom I had first interviewed in 2001) unexpectedly attended an early presentation of this article at the Fundação Getulio Vargas in March 2005. After the talk, he came forward to acknowledge the import threat, arguing that "foreign cement when dumped on our markets messes it up."

[14] These were several hundred thousand neighborhood stores to whom the industry typically delivered directly and who then sold on to small-scale consumers. In aggregate, these local stores accounted for 77% of purchases from producers in 1999, compared with only 11% by ready-mix concrete firms. Thus, four fifths of Brazilian cement was sold in bags rather than in bulk.

[15] The choice of W^I was influenced by a cement trader's cost spreadsheet that I had access to.

TABLE 1 The Explanatory Power of Imports Cost-Shifters and Domestic Cost-Shifters

Projection of Domestic Cement Prices onto Imports or Domestic Cost-Shifters						
Cost-shifters	**Imports**		**Domestic**		**Imp. & Dom.**	
Number of observations	3264		3264		3264	
Number of panel units	17 states, excl jungle		17 states, excl jungle		17 states, excl jungle	
Number of mo periods	192: 01/91–12/06		192: 01/91–12/06		192: 01/91–12/06	
	Coef.	SE	Coef.	SE	Coef.	SE
Imports cost-shifters, W^1						
Exchange rate (R$/US$)	4.93	(0.12)***			3.98	(0.12)***
International oil (WTI)	0.21	(0.02)***			0.12	(0.03)***
Maritime bulk freight (BDI)	0.11	(0.02)***			0.11	(0.02)***
Distance port × diesel	−0.50	(0.49)			−3.69	(0.58)***
Import finance	0.26	(0.02)***			0.16	(0.01)***
Immediate postdevaluation	−2.54	(0.15)***			−1.72	(0.13)***
Domestic cost-shifters, W						
Fuel oil (× fuel use)			2.27	(0.09)***	1.04	(0.09)***
Coal (× fuel use)			0.90	(0.08)***	0.42	(0.06)***
Electricity			0.27	(0.05)***	0.08	(0.04)*
Manufacturing wages			−9.82	(0.42)***	−2.45	(0.39)***
Distance plant × diesel			0.14	(0.09)	0.17	(0.07)**
Size of sourcing plants			−1.21	(0.22)***	−1.84	(0.17)***
Age of sourcing plants			−0.03	(0.13)	−0.06	(0.09)
Common cost-shifters, W or W^1						
Price controls 01/91–10/91	−2.18	(0.17)***	−6.02	(0.28)***	−3.76	(0.25)***
Intercept	Market fixed effects		Market fixed effects		Market fixed effects	
R^2	0.71		0.58		0.77	
R^2 excl market fixed effects	0.68		0.56		0.74	
R^2 fitted residuals regressed on the other set of cost-shifters	0.12		0.26			

Notes: Dependent variable is cement price (in constant December 1999 Reals per bag). Estimated through ordinary least squares (OLS). The spatial sourcing of each market (state) in the period 2004–2006 is unavailable, and thus I assume it to be the same as that in 2003. Standard errors are heteroskedasticity and autocorrelation robust (Newey-West 1 lag). Results are robust to dropping months between 2004 and 2006, or to adding the sparsely populated jungle states.
***Estimate that is significant(ly different from zero) at the 1% level; **5% level; *10% level.

(i) the exchange rate: the price of 1 U.S. dollar in (constant Brazil GPI) R$; (ii) the world price of oil (proxied by the West Texas Intermediate converted to R$ using the exchange rate); (iii) the price of maritime bulk freight (the Baltic Dry Index converted to R$); (iv) the price of inland transport: the domestic price of diesel oil for road transport, interacted with the local market's distance from the nearest port of entry; (v) the cost of import finance/bank letter of credit (proxied by the real yield on Brazilian Treasury bills); and (vi) a dummy variable to control for short periods (9 months each) following three large unexpected bouts of devaluation[16] during which the domestic industry would be wary that suddenly matching a higher limit price might be politically unpalatable. Variables specific to W are[17]: (i) and (ii) the domestic prices of two substitute kiln fuels—namely, fuel oil and coal—interacted with plant-specific fuel usage (e.g., cement plants in the south burned more coal, as this is where Brazil's coal mines were located); (iii) the domestic price of electricity, for grinding raw material and clinker; (iv) manufacturing

[16] Namely, the 1999 "Brazil crisis," the 2001 "Argentina crisis," and the 2002 presidential campaign.

[17] Plant-specific prices are weighted according to the plants' share of shipments to each market. Inland transport again considers diesel, as road is the predominant mode of shipment to buyers—about 90%.

wages; (v) the price of inland distribution: again the price of diesel, but now interacted with the market's average distance from the plants that supplied it; and (vi) and (vii) the (logarithms of the) average size and average age of the plants supplying the market. Finally, cost-shifters that are common to W^I and W are: (i) a dummy to account for price controls in the first ten months of 1991; and (ii) a set of market fixed effects.

The explanatory power of imports cost-shifters W^I is very high, with an R^2 of 71%—see the column marked "Imports." The exchange rate is highly significant, not only statistically but also economically: the coefficient (+4.93) multiplied by the mean exchange rate in the sample (1.72 R$/US$) accounts for 81% of the mean cement price (10.44 R$/bag). (And the lag in the exchange rate pass-through following large devaluation episodes such as the 62% devaluation of January 1999 suggests the presence of some dynamic effect.) The world oil price, the price of maritime freight, and the cost of import finance also have significantly positive effects. On the other hand, the transport-from-port interaction is negative (although insignificant, also economically), likely due to multicollinearity between this covariate and (i) the world oil price (correlated with the diesel oil price component of the interaction) and (ii) market fixed effects (which control for distance).[18] Of note, dropping the market fixed effects lowers the R^2 only slightly to 68%. This reflects the near uniformity of cement prices along coastal states, a fact that is consistent with the import price-ceiling hypothesis (I further comment on spatial price variation below).

The other price regressions indicate that the explanatory power of imports cost-shifters W^I survives when domestic cost-shifters W are controlled for. In the column "Domestic," the R^2 of a regression of cement prices p on W alone is a lower (but still quite high) 58%; taking the fitted residuals from this regression, $\hat{\varepsilon}_{p\perp W}$, and regressing them on W^I yields an R^2 of 26%. Notably, this owes chiefly to the exchange rate: regressing $\hat{\varepsilon}_{p\perp W}$ just on the exchange rate and a constant yields an R^2 as high as 18% (not reported). This orthogonal role of import prices can alternatively be seen in a regression of cement prices on both sets of covariates W^I and W: in the column marked "Imports & Domestic," coefficients on W^I are not too different from those in the first column, where only W^I were included as regressors (and the R^2 rises only slightly from 71% to 77%). In sum, Table 1 indicates that, despite the near absence of cement imports, import prices played an important role in determining the domestic price of cement.

Recent years nicely illustrate the role of the exchange rate. Between 2003 and 2006, the Real appreciated strongly against the U.S. dollar (due in part to growing world demand for Brazil's commodities): the average R$/US$ exchange rate in the final semester of 2006 was 44% lower (i.e., imports were cheaper) compared with that in the first semester of 2003. Over this same period, the domestic price of fuel oil fell 18% and the world price of oil, converted to R$ using the exchange rate, rose 15%.[19] Tellingly, domestic cement prices fell further than domestic energy prices, in the direction of the exchange rate: the average cement price in the 17 nonjungle states in the final semester of 2006 was 29% lower relative to that in the first semester of 2003 (8.48 vs. 11.94 R$/bag).

Two cross-sectional features of the price data are noteworthy. First, cement prices tend to increase in the distance from the nearest port of entry. This can best be seen by adding the 10 remaining interior "jungle" states (admittedly measurement-error prone—see below) to the above panel of 17 "populated" markets. For each of Brazil's 27 states, I calculate average annual cement prices over 16 years (1991–2006); I then regress these 432 market-year prices on distance from the nearest port of entry (in 1000 miles) and a set of year dummies. This yields a distance coefficient of 0.94 with standard error of 0.07. Second, cement prices tend to not vary along the coastal states (as noted above), despite large variation in local market structure. For example, the mean price (January 1991 to December 2006) in the coastal state of Santa Catarina (SC), where

[18] The transport-from-port coefficient becomes positive on dropping either (i) the world oil price (and significantly so), or (ii) the market fixed effects (although the effect is still insignificant, possibly because distance exhibits limited variation in the sample of consumer markets, as these are largely coastal).

[19] Notice that domestic energy prices, controlled by the government, did not move in line with the world oil price. (In US$ terms, the world oil price rose 107%, i.e., $(1 + .15)/(1 - .44) - 1 \simeq 1.07$.)

the mean one-firm concentration ratio (C_1) is a high 78%, is *similar* to that in the also coastal state of Rio de Janeiro (RJ), where C_1 is a lower 24%: a mean price of 10.75 R$/bag (standard deviation 2.16) in the former quasimonopolistic state against a mean price of 10.87 (SD 2.44) in the latter less-concentrated state. Of course, this feature could potentially be explained by a highly competitive domestic industry, but direct evidence of price-cost margins suggests otherwise.

Data: plant-to-market cement flows and (domestic) delivered marginal cost. I detail the sources and treatment of data in an online appendix. Here, I provide a brief overview and comment on actual (constructed) price-cost margins. The main data sources are the cement industry's trade association (SNIC)—notably quantity q—and the Brazilian Institute for Geography and Statistics (IBGE)—retail cement prices p and economic activity Y (either in construction and building, or aggregated across sectors of the economy[20]). In general, the data set consists of a panel, with the 27 Brazilian states as units over 156 months between January 1991 and December 2003. (Only for prices—p and cost-shifters W and W^1, as used in Table 1 above—does the data extend to December 2006.) In the subsequent structural analysis, I follow Table 1 and drop 10 sparsely populated northwestern states, due to possible measurement error. For comparison, despite being long established in Brazil, the global market research firm Nielsen does not audit jungle states due to their unusual geodemographic characteristics.[21] As specified in Section 2, the import price to end consumers p^I is treated as unobservable; the annual port-of-entry CIF series could be a proxy for its international component, but Brazil's minimal imports make it unreliable (this is why Figure 3 above depicts the U.S. CIF series).

The distinctive feature of the data is (domestic) supply. Most unusually, I observe *each plant's shipments to every state*, and thus the level at which each firm (given observed plant ownership) chose to supply (or not) each local market. Thanks to the simple fixed-coefficient production technology, I construct delivered marginal cost for every plant-market pair over time, which I later use to check my structural estimates. I combine plant-to-market shipments with observed (i) plant-level characteristics (e.g., location, capacity, technology, and age of kilns, and type of fuel used); (ii) engineering parameters; and (iii) local factor prices (e.g., fuel oil, coal, electricity, wages, and freight rates). Take, for example, the main component of plant marginal cost: to calculate the cost of kiln fuel for a specific plant, I multiply the marginal kiln's fuel economy (e.g., 750 kcal/kg of clinker, based on kiln characteristics) by the price of the specific fuel burned in the kiln (e.g., fuel oil). As for the plant-to-market specific freight cost, I do not directly observe freight rates paid by cement producers. But I proxy for these using data on equivalent freight rates for agricultural commodities collected (by the University of São Paulo) between 1997 and 2003 for thousands of different routes across Brazil. Soares and Caixeta Filho (1996) discuss how the transport of, say, soyabean and maize are close substitutes in the supply of cement freight. (Again, this is supported by several interviews.) As such, for any plant to any local market in any time period, I can accurately predict the marginal cost of producing and delivering a truckload of cement.

In calculating price-cost margins, I do not observe the cement producer's price to the reseller, only the retail price. But I can back out (net) producer prices by deducting (i) the reseller's cost and (ii) sales tax from retail prices. Based on interviews, I assume resellers were competitive. (The reseller's cost totaled about 12% of the retail price.) Based on tax legislation, I calculate sales tax, which was included in the producer's invoice to the reseller but collected by the producer (tax authorities naturally favored inspecting dozens of producer establishments rather than several hundred thousand resellers). Because I was able to directly obtain prices from a subset of producers, I use these to validate my backed-out producer prices, and the result is reassuring.

[20] Although cement is a critical input to construction, it accounts for only a small share of construction budgets. I will take construction activity to exogenously move the demand curve for cement (as does, for example, Syverson, 2004 for the downstream product concrete).

[21] The 10 states together accounted for 60% of Brazil's land but only 11% of its cement consumption. Note that on compiling consumption quantities I include sourcing from all plants in the country.

I find that price-cost margins for firm-market pairs where shipments *did* take place were very high: across producers, across states, and over time, the mean price-cost margin as a proportion of the producer price net of sales tax was 51% (the 10th percentile was 30% and the 90th percentile was 69%).[22] I subsequently discuss robustness checks of these margins. In sum, the industry wielded considerable market power.

□ **Demand: estimates can be rationalized by a price ceiling.** Let l index the 17 local consumer markets (nonjungle states) and $i = 0, \ldots, I$ index the cement plants that supply to them (with the aggregate fringe of foreign suppliers indexed by $i = 0$). The observed spatial sourcing can be described by a set of $(I + 1) \times 17$ matrices, one matrix for every time period t, where element q_{ilt} denotes the quantity of cement shipped by plant i for consumption in market l in that time period. Consumption is then $q_{lt} = \sum_{i=0}^{I} q_{ilt}$. The demand function (1) in each market l is $q_{lt} = D(p_{lt}, Y_{lt}, \varepsilon_{lt}^d; \alpha_l)$, where p_{lt} is the retail cement price and Y_{lt} are exogenous variables moving demand (i.e., construction activity). In the baseline demand specification, I estimate market-specific parameters α_l (thanks to the richness of the data) and adopt the linear form

$$q_{lt} = \alpha_{1l} + \alpha_{2l} Y_{lt} + \alpha_{3l} p_{lt} + \alpha_{4l} Y_{lt} p_{lt} + \varepsilon_{lt}^d, \tag{7}$$

where the interaction term $Y_{lt} p_{lt}$ allows the demand curve to rotate as exogenous demand varies. (I also allow the market-specific intercept α_{1l} to shift every quarter, to capture the seasonality of sales.) Employing 2SLS, I instrument for price (and its interaction) using a mixture of domestic and imports supply-shifters W_{lt} and W_{lt}^I (and their interaction with Y_{lt}), the identifying assumption being that they are uncorrelated with the unobserved component of demand.[23] I perform several robustness tests around these modelling choices, some of which are presented below.

Column I of Table 2 reports estimated price elasticities of demand in each of the 17 local markets for the baseline market-specific linear specification. For exposition, elasticities are evaluated at the mean value of covariates in the poststabilization phase, that is, the phase that covers about three quarters of the time period, from July 1994 on, during which Y_{lt} was higher in all states (more below). The poststabilization elasticity is negative for all 17 states, and significant at the 1% level in 16 states, varying from a minimum (in absolute) of -0.14 to a maximum of -0.64, with a mean of -0.38 and a standard deviation of 0.14. This is very low. Importantly, market demand is inelastic even in states where the supply of cement is quasimonopolistic: note, for example, the state of Santa Catarina (SC). Given the large number of parameters in these individual state regressions, only the coefficients on the interaction term are further reported. The interaction coefficient is negative in all 17 states, and significant at the 5% level in 14 states. There appears to be a common pattern across states, with demand becoming more elastic as demand curves shift out. Had elasticities in Table 2 been reported for the prestabilization phase, the mean elasticity across states would be an even lower -0.18. Thus, as prices in the economy stabilize and construction activity grows, the market price elasticity of demand doubles from about -0.2 to -0.4.

As for robustness, I fit alternative parametric specifications to the above baseline regression. Column II reports results for a (still market-specific) log-linear version of (7), taking quantity and price in logs. Results are similar, with the point estimate for the poststabilization elasticity ranging from -0.11 to -0.69, with a mean of -0.41 (and again significant at the 1% level in 16 out of 17 states). I can also impose theoretically plausible cross-market restrictions on the baseline specification, thus enhancing efficiency, as follows. Say that the "nature" of cement demand is similar across states such that the (linear) demand curve for state l swivels around a common price intercept—determined by a common highest-valuation buyer—according to the

[22] Salvo (2010a) presents summary statistics on shipments, prices, and costs, showing that there are many firm-market pairs for which shipments did *not* occur and yet margins were strikingly high.

[23] In the baseline specification, excluded instruments are namely domestic prices for fuel oil, coal, labor, and inland transport; the exchange rate; world prices for oil and maritime bulk freight; and the 1991 price control dummy. As mentioned, all prices are in constant R$.

TABLE 2 Market-Level Demand Estimates under Different Specifications

		Demand Estimates					
Specification		**I. Market-Specific Linear**		**II. Market-Specific Log Linear**		**III. Restricted Panel Linear**	
Number of observations		156		156		2652	
Number of panel units						17 states, excl jungle	
Number of mo periods		156: 01/91–12/03		156: 01/91–12/03		156: 01/91–12/03	
Elasticities (reported at the mean value of covariates in the poststabilization phase)							
	Cement consumption (kt, 1999)	Elasticity	SE	Elasticity	SE	Elasticity	SE
SP	11,723	−0.36	(0.06)***	−0.36	(0.06)***	−0.27	(0.03)***
MG	5,090	−0.51	(0.06)***	−0.50	(0.06)***	−0.32	(0.04)***
RJ	3,809	−0.49	(0.05)***	−0.50	(0.05)***	−0.35	(0.04)***
BA	2,461	−0.41	(0.07)***	−0.40	(0.07)***	−0.29	(0.03)***
PR	2,321	−0.25	(0.08)***	−0.26	(0.08)***	−0.37	(0.04)***
RS	2,221	−0.28	(0.07)***	−0.31	(0.08)***	−0.35	(0.04)***
SC	1,648	−0.15	(0.08)**	−0.25	(0.09)***	−0.17	(0.02)***
PE	1,225	−0.34	(0.06)***	−0.40	(0.06)***	−0.35	(0.04)***
CE	1,139	−0.50	(0.08)***	−0.57	(0.11)***	−0.44	(0.05)***
ES	837	−0.53	(0.05)***	−0.52	(0.06)***	−0.30	(0.03)***
MA	765	−0.58	(0.11)***	−0.57	(0.11)***	−0.15	(0.02)***
PB	565	−0.64	(0.08)***	−0.69	(0.07)***	−0.21	(0.02)***
RN	531	−0.28	(0.07)***	−0.30	(0.07)***	−0.28	(0.03)***
MS	454	−0.37	(0.06)***	−0.42	(0.07)***	−0.30	(0.03)***
AL	384	−0.28	(0.05)***	−0.35	(0.07)***	−0.22	(0.03)***
PI	379	−0.40	(0.09)***	−0.40	(0.09)***	−0.21	(0.02)***
SE	282	−0.14	(0.05)***	−0.11	(0.07)	−0.30	(0.03)***
Coefficients (for brevity, only interaction coefficient is reported for market-specific regressions)							
		Coef.	SE	Coef.	SE	Coef.	SE
Intercept		Market-specific		Market-specific		Restricted to Zero	
Construction activity, *Y*		Market-specific		Market-specific		32.31	(1.00)***
Cement price, *P* or ln *P*		Market-specific		Market-specific		Restricted to Zero	
Interaction, *YP* or *Y*ln *P*		Market-specific		Market-specific		−0.77	(0.09)***
SP		−5.48	(1.24)***	−0.04	(0.02)**		
MG		−4.61	(0.68)***	−0.08	(0.02)***		
RJ		−8.28	(1.15)***	−0.27	(0.05)***		
BA		−6.06	(0.95)***	−0.31	(0.07)***		
PR		−1.77	(0.88)**	−0.06	(0.05)		
RS		−1.67	(1.01)*	−0.06	(0.07)		
SC		−4.37	(2.52)*	−0.54	(0.27)**		
PE		−2.46	(0.94)***	−0.24	(0.14)*		
CE		−4.05	(0.66)***	−0.63	(0.12)***		
ES		−6.49	(1.20)***	−0.49	(0.22)**		
MA		−9.86	(2.16)***	−2.42	(0.67)***		
PB		−13.61	(2.08)***	−3.32	(0.61)***		
RN		−4.44	(1.03)***	−0.78	(0.44)*		
MS		−2.52	(1.17)**	−0.18	(0.41)		
AL		−4.26	(3.12)	0.35	(1.36)		
PI		−3.66	(0.92)***	1.35	(0.91)		
SE		−8.04	(2.33)***	−2.84	(1.21)**		

Notes: Results by state are in descending order of market size. Dependent variable is cement consumption (or its log, in the log-linear specification). Estimated through 2SLS. Excluded instruments are supply-shifters (domestic and imports) and their interaction with *Y*. Standard errors are heteroskedasticity and autocorrelation robust (Newey-West 1 lag). Significance levels: ***1%; **5%; *10%. Seasonality is captured through three quarterly dummies (columns I and II) or embedded directly in *Y* (column III).

state's construction activity (or population) Y_{lt}. As demand changes multiplicatively with Y_{lt}, that is, $p_{lt} = \tilde{\alpha}_1 + \tilde{\alpha}_2 Y_{lt}^{-1} q_{lt}$, this "restricted panel" demand function becomes

$$q_{lt} = \bar{\alpha}_1 Y_{lt} + \bar{\alpha}_2 Y_{lt} p_{lt} + \varepsilon_{lt}^d,$$

where the quantity intercept and the coefficient on the price-level term have been restricted to zero. Intuitively, the inverse demand curve $p(q)$ for a local market l that is twice the size of market l' would have the same price intercept but half the slope. As column III indicates, elasticities (all precisely estimated) remain of the same order of magnitude, ranging from -0.15 to -0.44.

A plausible interpretation. In sum, low market elasticities are systematic across local markets, including states where supply is highly concentrated. The (inverse) demand curve for cement is well known to be steep, owing to the product's typically low share of construction budgets and the absence of close substitutes (except in highway construction, where asphalt is a substitute). But although necessary, a steep demand curve is not sufficient to explain why demand is inelastic *at the equilibrium*. Why does an industry facing such inelastic demand not cut output to raise prices to a point where demand is more elastic? One possibility is that there is tough competition among incumbents such that the residual demand each *individual firm* faces is elastic. For example, low seller concentration might imply that any firm internalizes only a small fraction of the aggregate benefit (of the large price rise) that would result from a (small) reduction in output; in equilibrium, the price would then remain at a level consistent with aggregate demand being inelastic. I do allow for this domestic competition hypothesis when subsequently estimating supply, but one should note already that such a hypothesis is hard to square with observed concentration—C_1 averages 61% across states over 1991–1999—and observed price-cost margins—about 50% of producer prices. More plausible in light of the institutional setting and the price variation discussed previously is the hypothesis that a price limit binds on the domestic industry. By this hypothesis, whereas market demand is inelastic in equilibrium, the residual demand which the domestic industry faces at the price ceiling posed by high-cost imports is highly elastic.[24] Attempts by the domestic industry, already enjoying a large price-cost margin, to raise prices above this ceiling would only invite foreign entry.[25]

□ Supply: imports matter.

Imposing autarky. Mirroring Section 2, I first take the view of a researcher who assumes away a price-constraining role for imports (and who, of course, does not observe marginal cost). Table 3 reports 2SLS estimates of a (market-level) autarkic pricing equation (2), imposing the moment condition $E(Y'\varepsilon^s) = 0$ on

$$p_{lt} = -\theta_l q_{lt} \widehat{\partial p/\partial q}_{lt} + W_{lt}\beta + \varepsilon_{lt}^s, \tag{8}$$

where market marginal cost $W_{lt}\beta$ is flat in quantity q_{lt} and linear in local cost-shifters W_{lt}, and the coefficient θ_l on the markup term—which parameterizes specific nested oligopoly models—varies across states. (I present this specification for simplicity: results are robust to specifying the equation at the firm and market level, allowing marginal cost to vary in output, and letting θ_l vary over time, such as upon stabilization.) As the curvature of demand is estimated in an earlier step, rather than known, I compute bootstrapped standard errors.

Column A conditions on the baseline (market-specific) linear demand specification, and considers the full set of domestic cost-shifters W_{lt} included in the earlier exploratory exercise of

[24] It happens that the structural parameters of the DGP are such that market demand is inelastic at the ceiling. Other studies of cement have found low market price elasticities of demand, for example, -0.46 for Norway in Röller and Steen (2006) and -0.81 for the United States in Jans and Rosenbaum (1996).

[25] A third hypothesis hinges on a very special class of models of spatial competition, à la Hotelling-Salop, where a firm sets only a "mill" price and there is full pass-through of transport costs over space. This restrictive pricing scheme then ensures that a low *market* price elasticity of demand does not translate into a low price elasticity of demand faced by the *firm*.

TABLE 3 Estimation of a Standard Pricing Equation, without Modelling the Foreign Entry Threat, under Different Specifications

Supply Estimates, Import Constraint Unmodelled						
Supply specif. (domestic) conditional on demand	**A. Full *W* Col. I, Table 2 (Lin)**		**B. Full *W* Col. II, Table 2 (Log)**		**C. Reduced *W* Col. I, Table 2 (Lin)**	
Number of observations	2652		2652		2652	
Number of panel units	17 states, excl jungle		17 states, excl jungle		17 states, excl jungle	
Number of mo periods	156: 01/91–12/03		156: 01/91–12/03		156: 01/91–12/03	
	Coef.	Btstrp SE	Coef.	Btstrp SE	Coef.	Btstrp SE
Markup term, $q\partial p/\partial q$						
SP	0.02	(0.01)	0.03	(0.02)	0.02	(0.01)**
MG	0.03	(0.02)	0.04	(0.02)**	0.01	(0.01)
RJ	−0.01	(0.03)	0.01	(0.02)	−0.02	(0.03)
BA	−0.01	(0.01)	0.01	(0.01)	−0.02	(0.02)
PR	−0.04	(0.14)	0.00	(0.05)	0.01	(0.02)
RS	−0.08	(0.10)	−0.03	(0.04)	−0.07	(0.02)***
SC	0.01	(0.01)	0.00	(0.01)	0.00	(0.01)
PE	−0.01	(0.05)	0.02	(0.05)	0.04	(0.02)**
CE	−0.02	(0.01)*	0.00	(0.02)	−0.01	(0.02)
ES	0.01	(0.02)	0.08	(0.04)*	−0.06	(0.03)*
MA	0.01	(0.03)	0.01	(0.02)	0.01	(0.03)
PB	0.00	(0.02)	0.02	(0.02)	−0.01	(0.03)
RN	0.00	(0.01)	0.01	(0.03)	−0.01	(0.01)
MS	0.12	(0.08)	0.11	(0.07)	0.02	(0.01)*
AL	0.00	(0.08)	0.22	(0.09)**	0.00	(0.01)
PI	−0.02	(0.04)	−0.07	(0.07)	−0.02	(0.01)
SE	−0.01	(0.01)	0.00	(0.02)	−0.02	(0.02)
Domestic cost-shifters, W						
Fuel oil (× fuel use)	2.39	(0.19)***	1.93	(0.22)***	2.36	(0.16)***
Coal (× fuel use)	1.41	(0.17)***	1.34	(0.20)***	1.58	(0.21)***
Electricity	0.31	(0.07)***	0.32	(0.07)***	0.38	(0.12)***
Manufacturing wages	−5.53	(1.09)***	−4.19	(1.31)***		
Distance plant × diesel	0.61	(0.21)***	0.78	(0.24)***	0.22	(0.16)
Size of sourcing plants	−2.38	(0.51)***	−1.58	(0.58)***	−2.60	(0.38)***
Age of sourcing plants	−0.09	(0.30)	0.04	(0.35)		
Price controls 01/91–10/91	−4.95	(0.55)***	−4.31	(0.63)***	−2.13	(0.55)***
Intercept	Market fixed effects		Market fixed effects		7.55	(1.23)***

Notes: Dependent variable is cement price. Estimated through 2SLS, conditional on specified demand estimates of Table 2. Excluded instruments are demand-shifters (construction activity *Y*). Standard errors are estimated through bootstrapping, with 1000 replications, to account for demand estimation in the earlier step. Significance levels: ***1%; **5%; *10%.

Table 1 (recall column "Domestic," where cement prices p_{lt} were projected on W_{lt}). The present cost coefficients $\hat{\beta}$ are somewhat similar to those reported earlier, suggesting that the markup term estimated here contributes little to explaining p_{lt}: indeed, the markup parameters $\hat{\theta}_l$ are not statistically different from zero (perfect competition). The $\hat{\theta}_l$ are also quite precisely estimated.[26] By taking the observed (mean over t of the) reciprocal number of firms selling in a given state, the researcher might also test the null hypothesis of the industry behaving à la (static symmetric) Cournot toward that market, in which case he would reject Cournot in favor of competition in all but 2 out of 17 states (5% significance level; one-sided test results are not reported in Table 3359; only Paraná (PR) and Mato Grosso do Sul (MS) would not reject). For perspective, whereas an average of 6 firms sell in each market, a $\hat{\theta}_l$ of 0.01 corresponds to the equilibrium price-cost

[26] Precise (and downward-biased) estimates were also obtained in the Monte Carlo experiments mentioned earlier.

margin of a 100-firm Cournot industry which, at the mean (poststabilization) demand elasticity of -0.4 would amount to a margin of only $0.01/0.4 = 2.5\%$ of price (recall (3)). Alternatively, the researcher might estimate price-cost margins from $p_{lt} - W_{lt}\hat{\beta}$, obtaining tight 95% bootstrapped C.I.s that hover around zero, suggesting domestic competition. The researcher might find some reassurance in the significantly positive coefficients on most local factor prices (bar wages) and the significantly negative effect of plant size on cement prices.

Looking across Table 3, column A estimates are quite robust to (i) demand functional form: column B conditions on log-linear demand; and (ii) reducing the set of W_{lt}: column C's parsimonious specification drops wages (treating them as a fixed cost, considering a mostly supervisory role for labor; e.g., Norman, 1979), plant age (it is insignificant), and the set of market fixed effects (prices are near uniform along the 17 nonjungle states).

Relative to the near-zero C.I.s for the estimated price-cost margins, recall that the "actual" price-cost margins I constructed come in much higher.[27] One might ask whether my actual measure of the price-cost margin is accurate. When constructing "true" marginal cost, in several instances I chose to err on the side of caution, somewhat overstating cost and thus *understating* the true margin. For example, I assume that cement is composed of 100% clinker—by far the most expensive (energy-intensive) ingredient—when in reality four fifths of the industry's output was blended "type 2" cement with a lower 70%–80% clinker content. To provide another example, cement producers were among the country's largest industrial buyers, whereas the factor prices I observe are those paid by average buyers, not reflecting quantity discounts. I describe other examples of caution in the online appendix. The online appendix also details two reassuring tests of constructed marginal cost, based, respectively, on unusually disaggregate accounting data reported by one producer and on census-like data collected by the IBGE. In sum, direct evidence suggests that true price-cost margins were well above zero.[28]

Modelling the import threat. I now incorporate the price ceiling, in the expectation that allowing for a latent import threat will (i) reject the autarky assumption, and (ii) uncover significantly positive price-cost margins. Following Section 2, Table 4 reports ML estimates of a mixture model (4) based on the specification

$$p = \min\left(-\theta_l q_{lt}\widehat{\partial p/\partial q}_{lt} + W_{lt}\beta + \varepsilon^s_{lt},\, W^I_{lt}\beta^I + \varepsilon^I_{lt}\right), \tag{9}$$

where the delivered marginal cost of imports $W^I_{lt}\beta^I$ is linear in imports cost-shifters W^I_{lt}. I take as starting values the 2SLS estimates of the stand-alone domestic and imports equations. Given the estimation of demand in an earlier step, I again calculate bootstrapped standard errors. The columns of Table 4 correspond to the columns of Table 3—for example, column A again considers the baseline linear demand specification and the full set of domestic cost-shifters W_{lt}—now including the full set of imports cost-shifters W^I_{lt} from the exploratory exercise of Table 1. As results are fairly robust across columns (i.e., log-linear demand in column B and a parsimonious specification for W_{lt} in column C), the following comments center on column A estimates.

Perhaps surprisingly, the domestic supply estimates do not change significantly now that the latent role of imports is controlled for. Compared with column A of Table 3, the estimated markup parameters $\hat{\theta}_l$ tend to increase, but not significantly—a mean across the 17 states of 0.04 compared with 0.00 previously, and they tend to be less precisely estimated. As before, one

[27] In a previous version of this article, a figure illustrated the actual price-cost margins against C.I.s for estimated price-cost margins on sales to each of the three largest markets—São Paulo (SP), Minas Gerais (MG), and Rio de Janeiro (RJ). In the figure, tight C.I.s lay everywhere below the actual measures, with a bias of 3–4 R$ per bag averaging 57% of net producer prices.

[28] In constructing *marginal* cost I did not include the cost of capital, given the pervasive idle capacity. However, capital costs would amount to no more than half the constructed margin (i.e., 2 R$/bag).

TABLE 4 Estimation of Supply Allowing for the Foreign Entry Threat, under Different Specifications

	Supply Estimates, Import Constraint Modelled					
Supply specif. (domestic) conditional on demand	**A. Full *W* Col. I, Table 2 (Lin)**		**B. Full *W* Col. II, Table 2 (Log)**		**C. Reduced *W* Col. I, Table 2 (Lin)**	
Number of observations	2652		2652		2652	
Number of panel units	17 states, excl jungle		17 states, excl jungle		17 states, excl jungle	
Number of mo periods	156: 01/91–12/03		156: 01/91–12/03		156: 01/91–12/03	
	Coef.	Btstrp SE	Coef.	Btstrp SE	Coef.	Btstrp SE
Domestic supply parameters						
Standard Deviation of error (σ_s)	1.24	(0.22)***	1.32	(0.25)***	1.81	(0.25)***
Markup term $q\partial p/\partial q$:						
SP	0.11	(0.15)	0.04	(0.04)	0.03	(0.03)
MG	0.07	(0.10)	0.05	(0.05)	0.02	(0.03)
RJ	−0.01	(0.04)	0.01	(0.03)	−0.01	(0.02)
BA	−0.02	(0.02)	0.02	(0.03)	−0.01	(0.02)
PR	0.09	(0.14)	−0.02	(0.05)	0.03	(0.08)
RS	0.06	(0.09)	−0.02	(0.06)	−0.09	(0.07)
SC	0.01	(0.02)	0.00	(0.01)	0.01	(0.02)
PE	−0.03	(0.07)	0.02	(0.07)	0.09	(0.15)
CE	−0.03	(0.04)	0.00	(0.00)	0.01	(0.01)
ES	−0.01	(0.04)	0.04	(0.08)	−0.08	(0.06)
MA	0.04	(0.05)	0.00	(0.01)	0.01	(0.03)
PB	−0.01	(0.02)	0.04	(0.05)	−0.01	(0.02)
RN	−0.02	(0.03)	0.03	(0.05)	0.00	(0.01)
MS	0.43	(0.76)	0.07	(0.15)	0.01	(0.06)
AL	−0.03	(0.06)	0.07	(0.06)	0.00	(0.01)
PI	−0.03	(0.04)	0.03	(0.07)	0.02	(0.03)
SE	0.00	(0.00)	0.00	(0.00)	0.00	(0.00)
Domestic cost-shifters, *W*:						
Fuel oil (× fuel use)	2.73	(0.88)***	2.08	(0.74)***	2.44	(0.76)***
Coal (× fuel use)	1.17	(0.75)	1.34	(1.09)	1.54	(0.80)*
Electricity	0.80	(0.52)	1.13	(0.65)*	1.40	(0.66)**
Manufacturing wages	−3.67	(3.21)	−5.06	(2.62)*		
Distance plant × diesel	0.08	(0.23)	0.83	(1.12)	0.32	(0.27)
Size of sourcing plants	−9.33	(3.38)***	−3.14	(1.38)**	−4.40	(1.56)***
Age of sourcing plants	−0.10	(0.35)	−0.09	(0.52)		
Price controls 01/91–10/91	4.39	(17.54)	−2.23	(5.69)	−0.15	(1.03)
Intercept	Market fixed effects		Market fixed effects		6.16	(4.70)
Imports supply parameters						
Standard Deviation of error (σ_I)	1.19	(0.09)***	1.20	(0.09)***	1.35	(0.05)***
Cost-shifters, W^I:						
Exchange rate (R$/US$)	7.05	(0.39)***	6.90	(0.46)***	6.16	(0.52)***
International oil (WTI)	0.01	(0.06)	−0.05	(0.11)	0.12	(0.14)
Maritime bulk freight (BDI)	0.25	(0.11)**	0.16	(0.06)***	0.13	(0.06)**
Distance port × diesel	−0.99	(0.83)	−0.33	(0.74)	−1.07	(0.52)**
Import finance	0.09	(0.03)**	0.11	(0.03)***	0.18	(0.03)***
Immediate postdevaluation	−2.18	(0.32)***	−2.13	(0.32)***	−2.70	(0.26)***
Price controls 01/91–10/91	−3.79	(0.36)***	−3.62	(0.37)***	−3.23	(0.31)***
Intercept	Market fixed effects		Market fixed effects		−0.42	(0.52)
Log likelihood	−4042		−4143		−4490	
Sample mean Pr($\chi = 1$)	0.39		0.41		0.31	

Notes: Estimated through ML, conditional on specified demand estimates of Table 2. Standard errors are estimated through bootstrapping, with 1000 replications, to account for demand estimation in the earlier step. Starting values are 2SLS estimates of the two stand-alone equations. About 2% of replications did not converge and were excluded. Significance levels: ***1%; **5%; *10%.

cannot reject a null hypothesis of competition.[29] Standard errors on the domestic cost coefficients $\hat{\beta}$ are substantially higher. Estimated price-cost margins $p_{It} - W_{It}\hat{\beta}$ do not change in a meaningful way.[30]

On the other hand, the newly modelled imports supply side seems to fit the data well, yielding import cost coefficients $\hat{\beta}^I$ that are relatively precise.[31] The exchange-rate coefficient is larger (+7.05) than that obtained in the exploratory exercise of Table 1 (+3.98, column "Imports & Domestic"). As imports are less competitive when the R$/US$ rate is high (i.e., the Real is weak), the estimated implied probability that a market is unconstrained by imports (i.e., $\widehat{\Pr(\chi_{It} = 1)}$, recall Section 2) should correlate positively with the exchange rate: indeed, the two series are tightly correlated for each of the 17 states, with the correlation coefficient (over 150 post-stabilization months) ranging from 0.52 (Sergipe, SE) to 0.85 (SC) with a median of 0.77 (MS). Across all manner of specifications that I have taken to the data, the exchange rate towers over other W^I and all W in the significance—both statistical and economic—and stability of its coefficient.

What might explain the fact that on correcting for imports we do not uncover large price-cost margins? It is likely that model (9)—though richer than (8) in that it provides a role for imports—remains too simplistic to accommodate the data. If the behavior that governs (say) the unconstrained regime is not nested in the static domestic equation, this may contaminate supply estimates (Corts, 1999). For example, a real-world (and asymmetric) domestic cartel[32] may face incentive compatibility constraints on "monopoly" pricing, incur costs to switch between pricing regimes, and/or adopt focal pricing around an imports cost standard to facilitate coordination. Dealing with such dynamic issues would require more sophisticated models of behavior. Another possibility might be that there is an insufficient number of unconstrained markets in the sample, but a sample mean for $\widehat{\Pr(\chi_{It} = 1)}$ of about 30%–40% (see the last line of Table 4) would suggest otherwise.

Rejecting autarky. The Brazilian cement illustration shows that a researcher might have to make an allowance for the role of imports in determining price even when the quantity of imports is minimal. To make this point formally, step back for a moment and suppose that the researcher, observing virtually zero imports, has only *autarkic* oligopoly models (all theoretically plausible) in mind:

Model $a1$: Autarkic monopoly;
Model $a2$: Autarkic Cournot oligopoly; and
Model $a3$: Autarkic perfectly competitive industry.

The researcher fits each of these candidate models by ML and uses pairwise likelihood ratio (LR) Vuong (1989) tests (for nonnested hypotheses) to discriminate among them. (Conditioning on behavior to back out other supply characteristics such as cost is common practice in empirical IO, e.g., Bertrand in differentiated goods industries.) Table 5 reports test results: the bottom half uses the Schwarz (1978) criterion to adjust the LR for any differing degrees of freedom across each

[29] When taking Cournot as the null, Cournot is still rejected in favor of competition in as many as 12 states (only SP, MG, PR, Rio Grande do Sul (RS), and MS do not reject; test results again are not reported), compared to 15 states previously (i.e., all but PR and MS).

[30] In a previous version's plot of price-cost margins for the three largest markets (mentioned in note 27), point estimates $p_{It} - W_{It}\hat{\beta}$ that controlled for latent import competition lay everywhere above zero (and above the autarkic point estimates) only for São Paulo (SP), but even for this market the 95% C.I. was very wide, encompassing both zero and the actual (constructed) price-cost margins.

[31] The standard error (SE) on $\hat{\sigma}_I$ (0.09) is less than half that of its domestic equation counterpart $\hat{\sigma}_s$ (0.22). Compared to column "Imports & Domestic" of Table 1 (where p was projected on both W^I and W), SE's for $\hat{\beta}^I$ are about 3 times higher whereas SE's for $\hat{\beta}$ are about 10 times higher.

[32] Salvo (2010a) uses the rich plant-level supply data to uncover evidence of spatial collusion (and anecdotally cites a recent antitrust case that has been brought against the Brazilian cement industry on charges of explicit cartel formation).

TABLE 5 Pairwise Vuong Tests for Nonnested Models: Test Statistics and *p* Values

Different degrees of freedom not corrected	$a2$: autarkic Cournot	$a3$: autarkic competition	$i1$: integrated monopoly	$i2$: integrated Cournot
$a1$: autark. monopoly	−36.030 (0.000)	−49.605 (0.000)	−50.176 (0.000)	−50.176 (0.000)
$a2$: autark. Cournot		−35.975 (0.000)	−36.581 (0.000)	−36.581 (0.000)
$a3$: autark. competition			−4.856 (0.000)	−4.856 (0.000)
$i1$: integr. monopoly				−0.001 (0.500)
Correction according to Schwarz (1978)	**$a2$: autarkic Cournot**	**$a3$: autarkic competition**	**$i1$: integrated monopoly**	**$i2$: integrated Cournot**
$a1$: autark. monopoly	−36.030 (0.000)	−49.605 (0.000)	−49.756 (0.000)	−49.756 (0.000)
$a2$: autark. Cournot		−35.975 (0.000)	−36.221 (0.000)	−36.221 (0.000)
$a3$: autark. competition			−2.636 (0.004)	−2.636 (0.004)
$i1$: integr. monopoly				−0.001 (0.500)

Notes: Positive (negative) test statistic: model in row (model in column) displays better fit. *p* values are in parentheses.

tested pair of models, whereas the top half makes no correction.[33] Ignoring the columns (models) labeled $i1$ and $i2$ for the moment, having imposed autarky I obtain that perfect competition (model $a3$) wins "hands down" both over monopoly ($a1$) and over Cournot ($a2$)—this is expected from the θ-based Wald tests reported earlier. For example, the test statistic under the null hypothesis that autarkic Cournot and autarkic competition fit the data equally well is −36.0: given standard normal bidirectional cutoffs at 2.5% of +1.96 (the model in the row "Cournot" is favored) and −1.96 (the model in the column "competition" is favored), the statistic strongly rejects the null in favor of the alternative that autarkic competition fits the data better (with a p value of 0.000).

Now suppose that the researcher more generally enlarges the menu to include models of market power that allow for latent import competition—*integrated* models, such as[34]:

Model $i1$: integrated domestic monopoly; and
Model $i2$: integrated domestic Cournot oligopoly.

Whereas under the autarky assumption the perfectly competitive model $a3$ outperformed Cournot $a2$ (and monopoly $a1$), the selection tests in the last two columns of Table 5 indicate that either integrated model of market power, Cournot $i2$ or monopoly $i1$, outperforms (autarkic) perfect competition. For example, the null that autarkic competition and integrated Cournot have similar fits is strongly rejected in favor of the alternative that integrated Cournot fits the data better, with a p value of 0.004. Notice that in the setting at hand, I am unable to discriminate between integrated Cournot $i2$ and integrated monopoly $i1$; intuitively, conditional on either model, I estimate $\widehat{\Pr(\chi_{lt} = 1)}$ at or close to zero across markets, which would suggest that essentially all markets in the sample are constrained or that the unconstrained regime being specified does not adequately capture the data (as discussed above). However, the specification tests do conclusively (i) reject the hypothesis that Brazilian cement prices are determined by autarky in favor of a role for imports, and (ii) (in allowing a role for imports) preclude one from rejecting the presence of market power.

[33] For all competing models, I consider linear demand and the full set(s) of cost-shifters. (Parameter estimates are not reported for brevity.) Using Akaike's (1973) correction would yield the same results.

[34] An integrated perfectly competitive industry is not added to the menu, as this would require positive import quantities to be observed in every import-constrained market—import price below the competitive price—which is not the case in Brazilian cement. Recall note 4.

4. Conclusion

■ This article argues that overlooking the threat of imports can lead researchers to inadvertently reject domestic market power. As an empirical illustration, the Brazilian cement case provides mixed results. On the one hand, I uncover two intriguing features of the industry: (i) low equilibrium elasticities of demand that do not seem consistent with market power, yet direct measures of firm margins are too high to be consistent with competition; and (ii) despite minimal import quantities, import prices play an important role in determining the domestic cement price, so much so that the autarky hypothesis is formally rejected in favor of a role for imports. On the other hand, although these twin facts *can be rationalized* by the hypothesis that high-cost imports pose a price ceiling, the point I am as yet unable to show empirically is (iii) on correcting for the latent import threat we would find a monopolistic industry. Research that accounts for possible dynamics in the data may, in the future, provide (iii) as direct evidence of the price-ceiling mechanism.

The cement example shows us that one might have to make an allowance for the role of imports no matter whether they occur or not. More generally, the notion that trade can impact firm behavior even when trade flows are small appears increasingly relevant in a world where trade barriers are declining and yet some degree of local pricing power persists. In a sense, geographic markets may be coming closer together than meets the eye.

Appendix

This Appendix provides proofs and some simulation results.

□ **Proof of Proposition 1.** When $\chi_t = 1$, we have $p_t = p_t^* \leq p_t^I$ and $q_t = q_t^*$, and thus

$$(\chi_t = 1) \qquad \xi_t^s = p_t + \theta q_t \frac{\partial p(q_t, .)}{\partial q} - c(q_t, .) = p_t^* + \theta q_t^* \frac{\partial p(q_t^*, .)}{\partial q} - c(q_t^*, .) = \varepsilon_t^s. \tag{A1}$$

However, when $\chi_t = 0$, we have $p_t = p_t^I < p_t^*$ and $q_t = D(p_t^I, Y_t, \varepsilon_t^d; \alpha) > q_t^*$. From the second-order condition (SOC) (cf the first-order condition (FOC) (2)), we may write[35]

$$(\chi_t = 0) \qquad \xi_t^s = p_t + \theta q_t \frac{\partial p(q_t, .)}{\partial q} - c(q_t, .) < p_t^* + \theta q_t^* \frac{\partial p(q_t^*, .)}{\partial q} - c(q_t^*, .) = \varepsilon_t^s. \tag{A2}$$

Stacking up the full sample of observations $\chi_t \in \{0, 1\}$, and recalling that $E(Y'\varepsilon^s) = 0$ and $Y > 0$, it follows that

$$E(Y'\xi^s) < E(Y'\varepsilon^s) = 0.$$

To simplify exposition, denote $X_1 \equiv -q\partial p(q, .)/\partial q$. Now write marginal cost $c(q, W; \beta) \equiv X_2\beta$, where X_2 is an $N \times (K-1)$ matrix, each column corresponding to either a function of factor prices—label these exogenous terms X_2^{EXOG}—or a function of quantity and factor prices. (For example, X_2^{EXOG} corresponds to W in the case of flat marginal cost $c = W\beta$, and to W_1 in the case of linear marginal cost $c = W_1\beta_1 + W_2 q\beta_2$; from theory, note that each additive term is homogenous of degree one in factor prices.) Group the regressors of the estimated model into an $N \times K$ matrix, $X \equiv (X_1\ X_2)$, and the parameters to be estimated into a $K \times 1$ vector $\delta \equiv (\theta, \beta)$. The estimated model is then $p = X\delta + \xi^s$. Denote the matrix of instruments by $Z \equiv (Y\ X_2^{EXOG})$. From (A1) and (A2), noting that $E(X_2^{EXOG'}\varepsilon^s) = 0$ and $X_2^{EXOG} > 0$, it similarly follows that $E(X_2^{EXOG'}\xi^s) < E(X_2^{EXOG'}\varepsilon^s) = 0$. Thus,

$$E(Z'\xi^s) < 0. \tag{A3}$$

Assuming that the rank condition for identification holds (i.e., if Z has order $N \times J$, then $J \geq K$), the IV/2SLS estimator is then given by

$$\begin{aligned}\widehat{\delta} &= (X'Z(Z'Z)^{-1}Z'X)^{-1}X'Z(Z'Z)^{-1}Z'p \\ &= \delta + \left(\frac{1}{N}X'Z\left(\frac{1}{N}Z'Z\right)^{-1}\frac{1}{N}Z'X\right)^{-1}\frac{1}{N}X'Z\left(\frac{1}{N}Z'Z\right)^{-1}\frac{1}{N}Z'\xi^s.\end{aligned}$$

Applying the law of large numbers to each term along with Slutsky's theorem yields

$$\operatorname{plim}\widehat{\delta} = \delta + \left(E(X'Z)E^{-1}(Z'Z)E(Z'X)\right)^{-1}E(X'Z)E^{-1}(Z'Z)E(Z'\xi^s),$$

[35] To see this, note that (i) the left equality follows by construction of (5), (ii) the right equality follows from the FOC (2), and (iii) the SOC $\frac{\partial}{\partial q}(p(q, .) + \theta q\frac{\partial p(q,.)}{\partial q} - c(q, .) - \varepsilon^s) < 0$ and $q_t > q_t^*$ imply that

$$p_t + \theta q_t \frac{\partial p(q_t, .)}{\partial q} - c(q_t, .) - \varepsilon_t^s < p_t^* + \theta q_t^* \frac{\partial p(q_t^*, .)}{\partial q} - c(q_t^*, .) - \varepsilon_t^s.$$

where, in view of (A3) and the fact that $(E(X'Z)E^{-1}(Z'Z)E(Z'X))^{-1}E(X'Z)E^{-1}(Z'Z) \neq 0$, we have plim $\widehat{\delta} \neq \delta$.

□ **Monte Carlo experiments.** I report results for one Monte Carlo experiment among many that I have conducted under a wide range of demand and supply specifications. I motivate the setup by reference to a domestic monopolist that faces latent high-cost import competition. Demand is linear and domestic marginal cost is flat in output. The DGP is given by (4), and $c^I > c(D(c^I, .), .)$ holds for all realizations of the exogenous variables. I specify $S = 2000$ simulations, where each simulated data set (indexed by s) contains $N = 1000$ observations (indexed by t).[36]

□ **Experiment 1: linear demand and flat marginal cost monopoly.** In a given local market (for cement, say), demand is linear and is given by

$$q = D(p, Y, \varepsilon^d; \alpha) = \alpha_1 + \alpha_2 p + \alpha_3 Y_1 + \alpha_4 p Y_2 + \varepsilon^d,$$

where (i) $\alpha_1, \alpha_3 > 0$ and $\alpha_2 < 0$; (ii) observed exogenous variables Y_1, Y_2 are normally distributed on the positive real line, have diagonal covariance matrix, and $\alpha_2 + \alpha_4 Y_2 < 0$ for all realizations of Y_2; and (iii) $\varepsilon^d \sim N(0, \sigma_d^2)$. A domestic monopolist's marginal cost is flat in quantity

$$c(q, W; \beta) = c(W; \beta) = \beta_1 W_1 e + \beta_2 W_2,$$

where (i) $W_1 \sim N(\overline{W}_1, \sigma_{W_1}^2)$ is the price of a factor (oil, say), in US\$, traded on the world market; (ii) $e \sim N(1, \sigma_e^2)$ is the (real) exchange rate in LC\$/US\$ (LC is local currency); and (iii) $W_2 \sim N(\overline{W}_2, \sigma_{W_2}^2)$ is the price of a "local" factor (labor, say), set in the domestic market in LC\$. Imports can be supplied elastically at high cost

$$c^I(W^I; \beta^I) = \beta_1(1 + \beta_4)W_1 e + \beta_3 W_3 e$$

such that $c^I > c$ for all realizations, where (i) $\beta_4 \geq 0$ reflects the (energy) cost of international transport; and (ii) $W_3 = W_2/e + T$ is the "world" price of labor, in US\$, and $T \sim N(\overline{T}, \sigma_T^2)$ is the "trade cost" of labor. All cost coefficients β are positive and all cost covariates W, T are distributed over the positive real line. The errors of the DGP (4) are $\varepsilon^s \sim N(0, \sigma_s^2)$ and $\varepsilon^I \sim N(0, \sigma_I^2)$. In view of the domestic monopoly, $\theta = 1$.

Covariates and parameters. Complete the specification of the structural model by

$$Y_1 \sim N(20, 1^2), Y_2 \sim N(1, .1^2)$$
$$\alpha_1 = 10, \alpha_2 = -1, \alpha_3 = 1, \alpha_4 = -.2, \sigma_d = 1$$

on the demand side, and

$$W_1 \sim N(3, .3^2), W_2 \sim N(3, .3^2), e \sim N(1, .1^2), T \sim N(2, .2^2)$$
$$\beta_1 = 1, \beta_2 = 1, \beta_3 = 1, \beta_4 = 2.5, \sigma_s = 1, \sigma_I = 1$$

on the supply side. The reduced-form solution to the constraint-free system is then

$$p^*(1 + \theta) = -\theta\left(\frac{\alpha_1 + \alpha_3 Y_1 + \varepsilon^d}{\alpha_2 + \alpha_4 Y_2}\right) + \beta_1 W_1 e + \beta_2 W_2 + \varepsilon^s.$$

The proportion of constrained markets t, where $\min(p_t^*, p_t^I) = p_t^I$, happens to range from 45% to 56% over simulations s. (I vary this proportion in other experiments.)

Demand. Demand estimates using observed variables p, q are consistent—see Table A1.

Supply: imposing autarky. On the supply side, take the autarkic pricing equation (5) as the estimated model—see Table A2 for estimates. Whereas estimation with latent variables p^*, q^* would be consistent—there is no censoring by definition, as in Bresnahan (1982) or Porter (1983)—estimation with observed variables p, q understates the true degree of market power. The 95% C.I. for the elasticity-adjusted price-cost markup $-\eta(p - \widehat{c})/p$ lies below the actual realized (and sometimes constrained) markup $-\eta(p - c)/p$ in *all* of the $SN = 2 \times 10^6$ simulated markets. The downward bias is sizable. The estimated markup averages (i) 36% across unconstrained observations (against a realized markup of 98%), and (ii) 32% across constrained observations (against a realized markup of 77%).

Supply: modelling the import threat. Now implement the mixture model (4). Maximization of the log-likelihood function (6) consistently recovers the true supply parameters—see Table A3.[37]

[36] The online appendix reports results for three other experiments. In a second experiment, demand is log linear and domestic marginal cost is linear in output. In a third experiment, both demand and domestic marginal cost are linear, and the covariance matrix of the exogenous variables is no longer diagonal (e.g., $Corr(Y, W) > 0$). In a fourth experiment, demand is exponential and domestic marginal cost is linear. Results and the code are available upon request.

[37] In the experiment, W_2 and W_3 are imperfectly correlated, as I believe this to be more realistic. However, were W_2 and W_3 specified to be collinear, supply may still be recovered (due to the differential effect of demand on domestic and foreign supply).

TABLE A1 Experiment 1: Demand Estimates

Demand Estimates
2SLS estimation with full sample $\chi_t \in \{0, 1\}$ of:

	Observed Variables p, q	**Latent Variables** p^*, q^*	
Mean (standard deviation) over simulations s:			[Actual]
$\widehat{\alpha}_{1,s}$	10.01 (0.75)	9.87 (1.12)	$[\alpha_1 = 10]$
$\widehat{\alpha}_{2,s}$	−1.00 (0.04)	−0.98 (0.11)	$[\alpha_2 = -1]$
$\widehat{\alpha}_{3,s}$	1.00 (0.03)	0.99 (0.06)	$[\alpha_3 = 1]$
$\widehat{\alpha}_{4,s}$	−0.20 (0.02)	−0.20 (0.03)	$[\alpha_4 = -.2]$

Notes: Excluded variables instrumenting for p and pY_2 are W_1, W_2, W_3 and their interactions with Y_2. Similar estimates obtain with just identification (e.g., using W_1 and W_1Y_2 as instruments). Interacting instruments with e (say W_1e) yields similar estimates.

TABLE A2 Experiment 1: Supply Estimates without the Foreign Entry Threat

Supply Estimates, Import Constraint Unmodelled
2SLS estimation with full sample $\chi_t \in \{0, 1\}$ of:

	Observed Variables p, q	**Latent Variables** p^*, q^*	
Mean (standard deviation) over simulations s:			[Actual]
$\widehat{\theta}_s$	0.33 (0.02)	1.00 (0.03)	$[\theta = 1]$
$\widehat{\beta}_{1,s}$	2.78 (0.07)	1.00 (0.07)	$[\beta_1 = 1]$
$\widehat{\beta}_{2,s}$	1.00 (0.10)	1.00 (0.09)	$[\beta_2 = 1]$

Estimation with observed variables p, q: proportion of the SN markets such that the
95% C.I. for the estimated markup $-\eta\frac{p-\widehat{c}}{p}$ lies below the actual markup $-\eta\frac{p-c}{p}$: 100%.
Mean over s of means over $t \in s$ (standard deviation over s of means over $t \in s$), where t is...

...unconstrained	Estimate $-\eta\frac{p-\widehat{c}}{p}$:	[Actual $-\eta\frac{p-c}{p}$:	Estimate $\widehat{c}$:	[Actual c:
$(\chi_t = 1)$	0.36 (0.03)	0.98 (0.00)]	12.06 (0.24)	6.27 (0.02)]
...constrained	Estimate $-\eta\frac{p-\widehat{c}}{p}$:	[Actual $-\eta\frac{p-c}{p}$:	Estimate $\widehat{c}$:	[Actual c:
$(\chi_t = 0)$	0.32 (0.02)	0.77 (0.01)]	10.64 (0.23)	5.74 (0.02)]

Notes: Excluded variables instrumenting for the endogenous variable $-q\,\partial p(q,.)/\partial q = -q(\alpha_2 + \alpha_4 Y_2)^{-1}$ are Y_1, Y_2 (similar estimates obtain if $Y_1(\alpha_2 + \alpha_4 Y_2)^{-1}$, $Y_2(\alpha_2 + \alpha_4 Y_2)^{-1}$ are used). Similar estimates obtain if a constant is added to the estimated model. I assume that the market price elasticity of demand η is known, to abstract away from sampling error and focus on the supply inconsistency. Alternatively, using fitted estimates $\widehat{\eta}$ obtained in a first-stage demand estimation yields similar supply estimates.

TABLE A3 Experiment 1: Supply Estimates Allowing for the Foreign Entry Threat

Supply Estimates, Import Constraint Modelled
ML estimation with full sample $\chi_t \in \{0, 1\}$ of:

	Observed Variables p, q (with censoring condition accounted for)	
Mean (standard deviation) over simulations s:		[Actual]
$\widehat{\theta}_s$	1.00 (0.06)	$[\theta = 1]$
$\widehat{\beta}_{1,s}$	1.00 (0.13)	$[\beta_1 = 1]$
$\widehat{\beta}_{2,s}$	0.99 (0.14)	$[\beta_2 = 1]$
$\widehat{\beta}_{3,s}$	1.00 (0.08)	$[\beta_3 = 1]$
$\widehat{\beta}_{4,s}$	2.57 (0.53)	$[\beta_4 = 2.5]$
$\widehat{\sigma}_{s,s}$	1.00 (0.05)	$[\sigma_s = 1]$
$\widehat{\sigma}_{I,s}$	1.00 (0.04)	$[\sigma_I = 1]$

Notes: I again assume that demand (η) is known, to abstract away from sampling error and focus on the estimation of supply. Alternatively, using fitted estimates $\widehat{\eta}$ from a first stage yields very similar supply estimates. Estimates are robust to the choice of optimizer (either global or derivative-based search algorithms), optimization settings (e.g., termination tolerances), and starting values (selected from a fixed grid, a constant times the true parameters, or 2SLS estimates of the stand-alone domestic and imports equation components).

Raising the proportion of constrained markets even further. $\Pr(\chi_t = 1) \to 0$. Recall that in this experiment about half the markets are constrained. In the online appendix, I modify the setup to make the price ceiling bind further, sharply raising the proportion of constrained markets in the simulated industry to the 96%–99% range. On first imposing autarky, I obtain a mean $\widehat{\theta}$ across simulations of 0.01 (SD 0.01), against an actual $\theta = 1$, and the downward bias in the estimated markup is severe (e.g., $-\eta(p - \widehat{c})/p$ averages 1% across constrained observations, against a realized markup of 49%). I then estimate the true model (4). Due to the low number of unconstrained markets in each simulated sample, there is much variation in the estimated domestic supply parameters. I am unable to consistently recover domestic supply parameters. On the other hand, estimated imports cost parameters are consistent. Intuitively, variation in the domestic supply parameters around a large subregion of parameter space hardly impacts the log-likelihood function, as observed prices are largely driven by the imports pricing equation (and the demand curve).

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