Inferring Conduct under the Threat of Entry: The Case of the Brazilian Cement Industry

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Abstract

This paper demonstrates that when an industry faces potential entry and this threat of entry constrains pre-entry prices, the comparative statics of equilibrium will not identify supply. In such a setting, the identifying assumption behind the well-established technique of relying on exogenous demand perturbations to distinguish empirically between alternative hypotheses of firm behavior is shown to fail. The finding is highly relevant since (i) the use of the technique in both academic and policy circles is widespread, and (ii) the extent to which a variety of unobserved constraints restrain firms' ability to price, in a manner analogous to the threat of entry, is empirically important. The Brazilian cement industry, where the threat of imports restrains market outcomes, provides an empirical illustration. In particular, price-cost margins estimated using this established technique are biased heavily downwards, underestimating the degree of market power. Thanks to an unusually rich dataset, the paper further documents the supply pattern of a spatial cartel that is characterized by the tacit division of geographic markets.

Keywords: Market power; firm conduct; supply estimation; threat of entry; trade arbitrage; limit pricing; spatial cartel; market division; market sharing

JEL classification: D43, L13, L41, F14

1 Introduction

In the simple version of the static oligopoly model of profit maximization, firms set prices (or quantities) to solve unconstrained optimization problems, maximizing single-period revenues minus costs according to some assumption of firm conduct. The problem becomes more complex when industry characteristics, such as consumer stockpiling, network effects or learning by doing, introduce dynamic effects. These dynamic elements can be viewed as constraining firms' static pricing decisions, in the sense that they impinge on current market outcomes. Besides such industries whose characteristics naturally give rise to dynamic considerations, mounting empirical evidence documents other dynamic mechanisms, somewhat less observable *to the researcher*, by which firms may be constrained in their ability to set prices. One such instance is the *threat* of entry. This has received considerable theoretical attention but empirical studies are few, largely because of the unobserved nature of a threat. For example, Goolsbee and Syverson (2005) find that the threat of entry restrains pre-entry prices set by incumbents in the airline industry, while Genesove and Mullin (1998) suggest that the threat of imports explains the low prices set by the highly concentrated US sugar industry at the turn of the 20th century. In a similar vein, the threat of future price cap regulation has also been found to constrain current firm pricing in several industries, including oil (e.g. Erfle and McMillan 1990), electricity (Wolfram 1999), credit cards (Stango 2000) and pharmaceuticals (Ellison and Wolfram 2004).

This paper examines the performance of a well-established approach for estimating supply in a setting where unobserved threats constrain the ability of firms to set prices. Economists have long been concerned with the estimation of supply with a view to measuring, in the absence of cost data, the degree of market power exercised by firms, or their price-cost margins. Developed by Bresnahan (1982) and Lau (1982), the established approach hinges on the comparative statics of equilibrium to identify firm conduct and cost as specified in a static parametric pricing equation which follows from the firm's unconstrained optimization problem. Stated simply, the way equilibrium prices vary as demand conditions move exogenously reveals market power. Two polar examples provide intuition. In a competitive industry, firms set output at the point where price equals marginal cost. At the other extreme, a collusive industry, or a cartel, changes prices such that marginal revenue equals marginal cost. (I henceforth refer to the established approach as the *standard methodology*.) In contrast to this identification result, I demonstrate that in the more general setting where an industry's pricing is constrained by, say, the threat of entry, the standard methodology yields inconsistent estimates of market power, and may lead to a *downward* bias in particular. Intuitively, the threat of entry acts to constrain the ability of firms with market power to respond to exogenous demand shocks. Formally, the conventional identifying assumption of orthogonality be-

tween the error term of the standard pricing equation – which does not account for the constraint posed by the underlying threat – and the excluded exogenous variables which move the demand curve, does not hold.

To be clear, a researcher who overlooks the threat of entry or regulation when this threat has bite on a subset of the available market observations will, in a wide range of settings, underestimate market power, finding more competition when there is less. Importantly, the widespread use of the standard methodology over the past two decades, both in academia and in practice, suggests that this result is of high practical relevance.

I model an unobserved (and generally varying) price ceiling by reference to a simple limit-price model. A domestic oligopoly faces a competitive fringe of elastically-supplied high-cost imports. In equilibrium, no imports occur yet the threat of entry (imports) sets an upper limit on prices, equal to the delivered (marginal) cost of imports. In addition to providing a conceptual framework, this structural model paves the way for an empirical illustration. In the Brazilian cement industry, potential imports restrain market outcomes. The price ceiling imposed by imports binds at the equilibrium, such that no (or few) imports occur. I take on the role of the researcher who overlooks the underlying imports-arbitrage constraint and use the standard methodology to estimate conduct and cost. The supply *estimates* I obtain are indicative of domestic competition, with price-cost margins centered around zero. Now, the simple technology of the cement industry allows me to observe (construct) domestic marginal cost so that I can check these estimates. Cement is a given amount of limestone, a given amount of thermal energy to fire up the kiln, a given amount of electrical energy to grind the intermediate product, and in my data I observe the flow of cement from each plant to each local market, enabling me to calculate the cost of freight. The *true* (constrained) price-cost margins are far from competitive, amounting to around 50% of producer prices. Producers enjoy considerable market power despite the binding high-cost imports constraint[1]. This illustration confirms my theoretical proposition that the standard methodology in such constrained settings can yield biased supply estimates in the direction of more competition.

More generally, any constraint on the ability of firms with market power to set prices in response to changing demand conditions, will lead to a failure of the standard methodology's moment conditions for identification. There is a large body of game-theoretic work looking at the rationality of limit pricing. This work arose in response to the classic Chicago-School view that an incumbent's pre-entry pricing behavior should not be constrained if what is relevant to the entry decision is the post-entry price *and not the pre-entry one*. Countering this static view, price constraints in the entry deterrence

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[1]That the high price ceiling set by the high-cost imports binds is then a consequence of the steepness of the demand curve and this strong price discipline in the domestic industry.

literature are motivated, for example, by reference to a signaling game (prices reveal information on the incumbent's cost, the likelihood of predation, or the state of demand), or by reference to a real-options model with hysteresis ("it is easier to keep an entrant out than to drive him out"). With regard to the regulatory threat, Glazer and McMillan (1993) show how the threat of future intervention can restrain a monopolist's current price. The regulatory threat can also be cast in the form of an antitrust authority which exerts, in unobserved ways, downward pressure on an oligopoly's prices[2]. In fact, limit prices can be obtained in simple settings. One classic example is the price-setting game with heterogeneous firms, where the limit price is the (flat) marginal cost of the second most efficient firm (assuming this is lower than the monopoly price), in a manner analogous to the marginal cost of imports in this paper's model[3]. Indeed, many models in the international economics and macroeconomics literatures take the world price (plus trade cost) as a domestic limit price, under the implicit *contestability* (or *arbitrage*) assumption that the threat of entry is present in the form of opportunistic trade flows in a thick and well-organized international market, with low entry (and exit) costs. Of further relevance, macroeconomists have long studied how menu costs and other frictions (e.g. long-term contracts between sellers and buyers) lead to price rigidity ("stickiness"). A recent example is the work by Rotemberg (2004, 2005) on how the "threat of customer anger" restrains firms' pricing. Relatedly, a strand in the "behavioral" economics literature, starting with Kahneman, Knetsch and Thaler (1986), studies the relation between firms' prices and consumers' perceptions of fairness.

As for dynamic games of tacit collusion, there are reasons why the price responses of a cartel to demand perturbations may be muted. One reason may be the desire to avoid detection (Harrington 2004, 2005). Given the need to coordinate, another reason may be the adoption of a collusive focal price, such as the delivered cost of imports by a domestic cartel, or an existing legal price cap[4]. Further, in a model of collusion with stochastic demand shocks reminiscent of Rotemberg and Saloner (1986), Corts (1999) demonstrates that the standard methodology may perform poorly. Intuitively, this is the case when the sample contains sufficiently high realizations of demand such that the cartel's incentive constraint binds and the fully collusive price is not sustainable[5]. Indeed, though his

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[2]Antitrust authorities are typically fond of "barking" (or "jawboning") at industries perceived to have market power in the hope of restraining prices. To the extent that some pressure is indeed exerted – i.e. some of the *bark having bite* – this threat of antitrust enforcement provides another channel by which the standard methodology yields biased estimates of market power.

[3]Alternatively, one can view the marginal firm (or imports) as entrants with access to obsolete technology that have previously exited the industry (with Bertrand pricing between incumbents and entrants).

[4]Harris (1984) considers the Eastman-Stykolt (1966) hypothesis, where protection may facilitate coordination in a domestic oligopoly by establishing a focal point equal to the world price plus tariff. Relatedly, Knittel and Stango (2003) document the role of regulatory price ceilings in facilitating coordination in credit cards.

[5]One can show that this is another setting where demand shocks are not orthogonal to the errors in the standard pricing equation.

criticism of the standard methodology is illustrated through a specific dynamic game of collusion which may display anticyclical pricing, Corts' (1999) seminal paper carries a general message: that while the estimated conduct parameter is determined by the *marginal* responsiveness of equilibrium quantity (and thus price and the mark-up) to exogenous perturbations of demand, market power is defined by the *level* of the price-cost margin. By reference to a simpler limit-price model, I am able – for a wide range of settings – to (i) confirm the validity of Corts' inconsistency argument, (ii) sign the inconsistency, and (iii) empirically illustrate the (downward) inconsistency.

With further regard to the Brazilian cement industry, I use the richness of the microdata – the observability of marginal cost and, most unusually, *shipments from each plant to each local market* across the country – to document the supply pattern of a spatial cartel that is characterized by the tacit division of geographic markets. Given the need to control for the imports-arbitrage constraint that sets an aggregate output floor (equivalently, a price ceiling) at each local market, I design a test that uses firm-level supply data to place a tighter lower bound on the collusiveness of firm conduct than would be possible with market-level data alone. I show that firm-level behavior in Brazilian cement is considerably more collusive as benchmarked against a Cournot model, with a given firm holding back supply to certain local markets (relative to the benchmark) in exchange for its rivals giving it the upper hand in other local markets. Mechanically, the test endogenously selects firm-level supply decisions in (geographic and time) markets which can only be explained by behavior that is "more collusive than Cournot".

This paper thus makes four contributions. *First*, it demonstrates that the widely-used methodology for inferring conduct may lead to the underestimation of market power in industries where equilibrium prices are constrained in ways that are unobserved to the researcher[6]. Potential competition from imports – one of several channels by which such constraints may bind – alone suggests that this result is increasingly relevant in a world where trade barriers are being pulled down. *Second*, thanks to a unique dataset, the paper provides a rare example of a cartel's allocation of spatial markets, let alone an example of a current spatial cartel responding to import competition. The *third* contribution relates to the specific empirical setting. In a developing country such as Brazil, with its huge housing deficit and infrastructure needs, the importance of the cement industry cannot be overstated. A clear policy recommendation is to reduce the transaction costs of imports, thus curbing the substantial market power of domestic producers. That the margin of adjustment will be (reduced) domestic prices rather than

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[6]Equivalently, one can consider constraints that cannot easily be built – via a regime-switching mechanism – into the structural model. This may be due to the researcher's lack of understanding of how the constraint operates, even though he may be aware of its existence, or where conditions for identification of the switch (and therefore of the unconstrained observations) are too stringent. See footnote 15.

(increased) import quantities might assuage political resistance[7]. *Finally*, the Brazilian cement case adds to a short list of empirical studies showing that incumbents do respond to the *threat* of actions by other agents (such as entrants, consumers and authorities), not only to their *actual* actions. However, and of interest to the theory of contestable markets, the case also illustrates that finding that the threat of entry has a binding effect on an industry's prices should not lead one to assume away the presence of market power.

The plan of the paper is as follows. In Section 2, I develop the theoretical framework and address identification. In Section 3, I consider institutional aspects of the cement industry, and present the data. Section 4 presents the application. Finally, I conclude, reflecting on the methodological implications of this paper.

2 Theoretical framework: The standard methodology under the threat of entry

I begin by intuitively discussing why the standard methodology for inferring supply can lead to a downward bias in the estimated degree of market power when the threat of entry constrains market outcomes. Potential entry is modeled as a competitive fringe of foreign suppliers to the domestic oligopoly market. I then formally demonstrate the failure of the orthogonality condition.

The static setting is depicted in Figure 1. Importantly, I momentarily take marginal cost to be flat in quantity only for the sake of illustration; the standard methodology[8], as well as my formal results that follow later, allow marginal cost more generally to vary in output. The researcher observes market outcome E_1 and wishes to identify whether this outcome has been generated by either of two alternative hypotheses, say: a low-cost cartel (or monopoly), with cost $\underline{c}$, or a high-cost competitive industry, with cost $\overline{c}$. Then an exogenous shift (resp. rotation) to demand is observed and the equilibrium changes to either E_2^M or E_2^C (resp. changes to E_3^M or remains at $E_1 = E_3^C$) according to the hypothesis of firm behavior. Collusion is observationally distinct from competition because the response of prices to demand shocks is different: while firms with market power change prices to ensure that marginal revenue is equated to marginal cost, in a competitive industry price equals marginal cost.

Now modify the low-cost monopoly hypothesis by allowing potential entry to constrain prices. A domestic monopolist faces a competitive fringe of high-cost imports,

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[7]Such a recommendation contrasts with recent policy experience, where the cement industry successfully lobbied the government into enacting "antidumping" measures against foreign producers.

[8]When marginal cost varies in quantity, rotations of the demand curve identify conduct (Bresnahan 1982).

labeled I, with perfectly-elastic supply at marginal cost $c^I > \underline{c}$. The equilibrium is given by either of two situations (see Figure 2). If the marginal cost of imports c^I is lower than the monopoly price in the absence of imports, denoted p^M, the market price will equal c^I, the monopolist will supply the entire domestic market, yet the foreign fringe exerts downward pressure on price. Alternatively, if $c^I \geq p^M$, imports have no "bite" and the equilibrium price will be p^M, with the monopolist again supplying the entire market though in an unconstrained manner. The equilibrium price is thus

$$p = \begin{cases} c^I & \text{if} \quad p^M > c^I \\ p^M & \text{otherwise} \end{cases}$$

where $p^M = p(q^M)$, $p(q)$ is the inverse demand function and q^M is the quantity that equates the market's marginal revenue $MR(q)$ to the monopolist's marginal cost $\underline{c}$. Given the assumption that $c^I > \underline{c}$, entry does not occur[9]. At the limit price $p = c^I$, while the price elasticity of (residual) demand faced by the monopolist is infinitely high in absolute value, the market price elasticity of demand, $\eta(q) := \frac{\partial \ln q}{\partial \ln p(q)}$, is finite (and possibly inelastic, as illustrated in the figure). Around this constrained equilibrium, fluctuations in the marginal cost of imports, say due to fluctuations in the exchange rate, identify the market demand curve, since the kinked equilibrium moves up and down along the demand curve.

In this constrained setting, do fluctuations in the demand curve identify conduct? Figure 3 indicates that this is no longer the case. Again the researcher initially observes equilibrium E_1 and wishes to empirically distinguish between two alternative hypotheses: that of a low-cost cartel, with cost $\underline{c}$ *and* with imports constraining prices at the high marginal cost of imports $c^I = \bar{c}$, against that of a high-cost competitive industry, with cost $\bar{c}$ (and where the presence of imports becomes irrelevant). Here, prices do not respond differently to demand shocks according to the behavioral hypothesis. Under both alternative hypotheses, a demand shift moves the equilibrium to E_2, while a rotation of the demand curve around the equilibrium point E_1 leaves the equilibrium unchanged. Intuitively, the threat of imports constrains the ability of firms with market power to set marginal revenue equal to marginal cost. Put differently, equilibrium market price elasticities of demand are no longer informative since the equilibrium lies at the kink of the residual demand curve facing the domestic oligopoly.

In sum, there is no observable distinction between the hypothesis of a constrained low-cost cartel and the hypothesis of high-cost competition. By overlooking the constraining

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[9]Clearly, when $p^M > c^I$, the result of imports commanding zero sales rests on the assumption of perfectly-elastic supply from the foreign fringe. The assumption serves only to make the paper's methodological point regarding the presence of unobserved constraints. To the extent that the marginal cost of the fringe is upward-sloping, by observing entry the researcher can naturally build the presence of the fringe into the structural model (as, for example, Suslow 1986 does).

effect of entry and misspecifying the structural model to capture the static setting of Figure 1, a researcher would inadvertently take the lack of price variation to exogenous movements in demand as evidence to reject collusion.

Estimation of a static pricing equation   The empirical literature on conduct typically assumes that 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 (aggregate, say) 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 $p(q, .) = D^{-1}(p, .)$ is inverse demand, $c(q, .)$ is marginal cost, Y and W are respectively exogenous demand and supply covariates that are observed, and the zero-mean error terms ε^d and ε^s respectively capture the exogenous components of demand and supply that are unobserved (and are orthogonal to Y and W). The structural researcher, "knowing" the functional forms of $D(p, .)$ and $c(q, .)$, wishes to estimate parameters α (demand), β (marginal cost), and θ (conduct)[10]. One reason why supply specification (2) may have become so popular is that it nests first-order conditions corresponding to the oligopoly models of monopoly or perfect collusion (where the firm internalizes the aggregate inframarginal revenue, so that $\theta = 1$) and perfect competition (where $\theta = 0$), among other models (e.g. symmetric Cournot, θ being the reciprocal of the number of firms in the industry)[11]. Pricing equation (2) can be rearranged to the familiar "elasticity-adjusted Lerner index" (or price-cost mark-up):

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

where η is the market price elasticity of demand (and where, for simplicity, arguments have been suppressed and ε^s has been subsumed into marginal cost).

Following (or simultaneous with) the estimation of demand (1) – which yields con-

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[10]Examples include Gollop and Roberts (1979), Porter (1983), Roberts (1984), Suslow (1986), Bresnahan (1987), Brander and Zhang (1990), Parker and Röller (1997), Kadiyali (1997), Wolfram (1999), Nevo (2001), Slade (2004) and Puller (2007). Genesove and Mullin (1998), Clay and Troesken (2003), and Kim and Knittel (2004) are concerned with testing the standard methodology – see below.

[11](2) can also be specified at the firm level, in which case a subscript f is added to c_f and θ_f. Averaging across these firm-level pricing equations (weighting or not by firms' shares) yields the industry-level pricing equation (2), where θ can be interpreted as "the average collusiveness of conduct" (Bresnahan 1989). A common alternative to the firm-level version of (2) replaces industry output q by firm output q_f in the inframarginal revenue term. θ_f then corresponds to dq/dq_f, which can be interpreted as a "conjectural variation" of aggregate output dq conjectured by firm f upon expanding its output by dq_f. For the purpose of my paper, I abstract from the discussion surrounding this interpretation (see, e.g., Reiss and Wolak 2005).

sistent estimates of α and thus of $\frac{\partial p(q,.)}{\partial q}$ – supply specification (2) is estimated by IV or GMM and the identifying assumption is

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

where excluded demand covariates Y (including "demand rotators" à la Bresnahan 1982) are assumed to be orthogonal to the error ε^s, serving as instruments for $q\frac{\partial p(q,.)}{\partial q}$ and the endogenous elements of $c(q,.)$.

This standard methodology fails in the presence of imports arbitrage since, to the extent that the threat of entry constrains a subset of market outcomes, fluctuations in the demand curve will be correlated with the error in the specified static pricing equation. Now denote the solution to the system given by the market demand function (1) and the *constraint-free* pricing equation (2) by p^* and q^*, i.e.

$$(p^*, q^*) = \arg \text{solve} \begin{cases} q = D(p, Y, \varepsilon^d; \alpha) \\ p = -\theta q \frac{\partial p(q, Y, \varepsilon^d; \alpha)}{\partial q} + c(q, W; \beta) + \varepsilon^s \end{cases}$$

Let the high (and exogenous – recall note 9) marginal cost of imports be

$$c^I = c^I(W^I; \beta^I)$$

where $c^I > c(D(c^I, .), .)$ for all realizations of the exogenous variables. (The condition ensures that imports do not occur in equilibrium; again, it serves only to make the paper's methodological point regarding the presence of unobserved pricing constraints. If imports did occur, and were thus observed by the researcher, we would be back in the world of, e.g., Suslow 1986.) The supply side of the true model – the Data Generating Process – is given by[12]

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

The structural researcher observes Y and W, and the limited dependent variables p and $q = D(p, Y, \varepsilon^d; \alpha)$, but is *unaware* of the imports-arbitrage constraint $p \leq c^I$ that is censoring the data. The endogenous variables p^* and q^* are latent. In contrast to the DGP, the estimated model is

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

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[12]It is clear from (4) that, ceteris paribus, the likelihood that the imports-arbitrage constraint binds, and thus $p = c^I$, is higher (i) the more collusive is conduct, i.e. the higher is θ; (ii) the steeper is the demand curve, i.e. the higher is $-q(\partial p(q, .)/\partial q)$; (iii) the higher is the domestic industry's marginal cost $c = c(q, .)$; and (iv) the lower is the marginal cost of imports c^I.

where the error of the (mis)specified pricing equation is denoted ξ^s. There are N observations. Label each observation i using the indicator function

$$\chi_i := \mathbf{1}[p_i^* \leq c_i^I] = \mathbf{1}[\varepsilon^s \leq \theta q^* \frac{\partial p(q^*, Y, \varepsilon^d; \alpha)}{\partial q} - c(q^*, W; \beta) + c^I(W^I; \beta^I)]$$

such that $\chi_i = 1$ corresponds to a market equilibrium which is unconstrained by the threat of entry and $\chi_i = 0$ corresponds to a constrained equilibrium. When $\chi_i = 1$, we have $p_i = p_i^* \leq c_i^I$ and $q_i = q_i^*$, and thus

$$(\chi_i = 1) \qquad \xi_i^s = p_i + \theta q_i \frac{\partial p(q_i, .)}{\partial q} - c(q_i, .) = p_i^* + \theta q_i^* \frac{\partial p(q_i^*, .)}{\partial q} - c(q_i^*, .) = \varepsilon_i^s \qquad (6)$$

However, when $\chi_i = 0$, we have $p_i = c_i^I < p_i^*$ and $q_i = D(c_i^I, Y_i, \varepsilon_i^d; \alpha) > q_i^*$. From the second-order condition (cf. the FOC (2)), we may write[13]

$$(\chi_i = 0) \qquad \xi_i^s = p_i + \theta q_i \frac{\partial p(q_i, .)}{\partial q} - c(q_i, .) < p_i^* + \theta q_i^* \frac{\partial p(q_i^*, .)}{\partial q} - c(q_i^*, .) = \varepsilon_i^s \qquad (7)$$

Clearly, the DGP is a generalization of the static model considered by Bresnahan (1982); this static model would correspond to a specific situation where $\Pr(\chi_i = 1) = \Pr(p_i^* \leq c_i^I)$ $= 1$. The theoretical specification (2) that underlies the estimated model (5) fails to adequately capture the supply decisions (4) of an industry with pricing power facing a price ceiling. This is summarized in the following proposition.

Proposition 1 (*Non-identification of the degree of market power*) *When the threat of entry constrains prices set by an industry with market power, the residual ξ^s in the standard pricing equation is negatively correlated with the excluded exogenous demand variables Y:*

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

Consequently, IV (or GMM) estimation using demand perturbations Y will yield inconsistent estimates of conduct and cost.

Proof. From (6) and (7), the errors of the estimated model are given by

$$\xi_{i|\chi_i=1}^s = \varepsilon_{i|\chi_i=1}^s \text{ and } \xi_{i|\chi_i=0}^s < \varepsilon_{i|\chi_i=0}^s \qquad (8)$$

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[13]To see this, note that the SOC $\frac{\partial}{\partial q}\left(p(q, .) + \theta q \frac{\partial p(q, .)}{\partial q} - c(q, .) - \varepsilon^s\right) < 0$ and $q_i > q_i^*$ imply that

$$
\begin{aligned}
p_i + \theta q_i \frac{\partial p(q_i, .)}{\partial q} - c(q_i, .) - \varepsilon_i^s \quad &< \quad p_i^* + \theta q_i^* \frac{\partial p(q_i^*, .)}{\partial q} - c(q_i^*, .) - \varepsilon_i^s \\
&= \quad 0 \quad \text{(by the FOC (2))} \\
&= \quad p_i + \theta q_i \frac{\partial p(q_i, .)}{\partial q} - c(q_i, .) - \xi_i^s \text{ (by construction of (5))}
\end{aligned}
$$

Stacking up the full sample of observations $\chi_i \in \{0, 1\}$, and maintaining the standard assumptions that (i) $E(Y'\varepsilon^s) = 0$ (i.e. the unobserved supply shock ε^s is orthogonal to the excluded exogenous demand variables Y), and (ii) $Y > 0$ (realizations of exogenous demand are positive), it follows that

$$E(Y'\xi^s) < E\left(Y'\varepsilon^s\right) = 0 \qquad (9)$$

To simplify exposition, denote $X_1 = -q\frac{\partial p(q,.)}{\partial q}$ and $c = c(q, W; \beta)$. Now write marginal cost as being linear in X_2, i.e. $c = X_2\beta$, where X_2 is an $N \times (K-1)$ matrix of observed variables, where each of the $(K-1)$ variables that enters additively is either exogenous (a function of factor prices) or endogenous (a function of quantity and factor prices, such that each additive term is h.d.1 in factor prices)[14]. Group the regressors of the estimated model into an $N \times K$ matrix, $X := (X_1 \ \ X_2)$, and the parameters to be estimated into a $K \times 1$ vector $\delta := (\theta, \beta)$. The estimated model is then $p = X\delta + \xi^s$. Denote as Z the matrix of instruments, containing the excluded exogenous demand variables Y and the exogenous variables of X_2, i.e. $Z = (Y \ \ X_2^{EXOG})$. From (8) and the assumptions that (i) $E(X_2^{EXOG\prime}\varepsilon^s) = 0$, and (ii) $X_2^{EXOG} > 0$ (realizations of the exogenous marginal cost covariates are positive), it similarly follows that $E(X_2^{EXOG\prime}\xi^s) < E\left(X_2^{EXOG\prime}\varepsilon^s\right) = 0$, and thus

$$E(Z'\xi^s) < 0 \qquad (10)$$

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

$$\begin{aligned} \hat{\delta} &= (X'Z(Z'Z)^{-1}Z'X)^{-1}X'Z(Z'Z)^{-1}Z'p \qquad (11)\\ &= \delta + (\frac{1}{N}X'Z(\frac{1}{N}Z'Z)^{-1}\frac{1}{N}Z'X)^{-1}\frac{1}{N}X'Z(\frac{1}{N}Z'Z)^{-1}\frac{1}{N}Z'\xi^s \end{aligned}$$

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

$$\text{plim } \hat{\delta} = \delta + \left(E\left(X'Z\right)E^{-1}\left(Z'Z\right)E\left(Z'X\right)\right)^{-1}E\left(X'Z\right)E^{-1}\left(Z'Z\right)E(Z'\xi^s) \qquad (12)$$

where, in view of (10) and the fact that $\left(E\left(X'Z\right)E^{-1}\left(Z'Z\right)E\left(Z'X\right)\right)^{-1}E\left(X'Z\right)E^{-1}\left(Z'Z\right) \neq 0$, we have plim $\hat{\delta} \neq \delta$. ∎

Corollary 1 *(Downward bias in the estimated degree of market power) For a wide class of demand $D(p, .)$ and marginal cost $c(q, .)$ specifications, the true degree of market power θ will be underestimated, i.e.*

$$plim \ \hat{\theta} < \theta$$

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[14]Examples include (i) linear cost (flat marginal cost), i.e. $c = W\beta$, (ii) quadratic cost, i.e. $c = W_1\beta_1 + W_2q\beta_2$, and (iii) Cobb-Douglas technology with parameters γ_k, i.e. $c = q^{\frac{1-\Sigma_k\gamma_k}{\Sigma_k\gamma_k}}\Pi_k(\frac{W_k}{\gamma_k})^{\frac{\gamma_k}{\Sigma_k\gamma_k}}$.

I have attempted to analytically demonstrate the negative sign of $\left(\text{plim } \hat{\theta} - \theta\right)$ for a wide class of specifications (by partitioning the matrix expression (12)). However, (as inspection of (12) indicates) the algebra quickly becomes very involved. For this reason, my strategy for proving Corollary 1 (included in the Appendix) is as follows. My point of departure is to analytically derive $\left(\text{plim } \hat{\theta} - \theta\right)$ for the (algebraically) simple case where (i) demand $D(p,.)$ is linear, (ii) marginal cost $c(q,.)$ is flat in output, (iii) the rank condition for identification just holds and $J = K = 2$, and (iv) the exogenous demand and supply covariates Y and W are independent. I then proceed to relax each of the assumptions of this simple case in Montecarlo experiments where (i) demand is concave or convex (e.g. exponential), (ii) marginal cost varies in output (e.g. linearly increasing), (iii) there is a large number of supply covariates (and there may be overidentification), and (iv) the covariance matrix of the exogenous variables is not diagonal (e.g. $Cov(Y, W) > 0$). For every set of simulated data $s = 1, ..., S$ (where $S = 2000$ say) in each one of the Montecarlo experiments I have designed, under a wide range of assumptions, *the conduct parameter estimated from the full sample* (of constrained and unconstrained observations $\chi_i \in \{0, 1\}$) *lies below the true conduct parameter (in a statistically significant sense).* That is, each one across a wide class of experiments yields:

$$\hat{\theta}_s^{full\ sample} < \theta,\ \forall s$$

Proposition 1 and its corollary have clear implications for empirical work. Consider an industry where firms have market power ($\theta > 0$) and the threat of entry constrains prices in equilibrium for at least a subset of the data (i.e. $\Pr(\chi_i = 0) > 0$ in the available sample). Suppose a researcher fails to realize this price-restraining effect and runs specification (5) on the data, thinking that the DGP is (2), when it is actually (4)[15]. Thinking that he is imposing $E(Y'\varepsilon^s) = 0$, when in fact he is incorrectly imposing $E(Y'\xi^s) = 0$, the researcher would obtain inconsistent estimates of conduct and cost. For a wide range of demand and supply situations, the estimated conduct parameter $\hat{\theta}$ would lie below the "true" metric $\theta = \eta\frac{p-c}{p}$, as defined in (3), underestimating the degree of market power. Intuitively, since the response of prices to demand shocks is

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[15]Note that were the researcher *aware* of the price-restraining effect of entry on a subset of the data, *and* were able to "separate the wheat from the chaff", in principle he could implement the standard methodology using the unconstrained outcomes only (i.e. modeling the truncated conditional distribution $f\left(\varepsilon_i^s \mid \chi_i = 1, Y_i, \varepsilon_i^d, W_i, W_i^I; \alpha, \beta, \theta, \beta^I\right)$), or he could use switching regression techniques (assuming, of course, that *not all observations are constrained*). Typically, however, (i) the constraining effect of potential entry may be overlooked since entry does not occur in equilibrium; (ii) the level at which the constraint binds fluctuates and is unobserved (e.g. in the example of footnote 2, the level at which the antitrust *bark* has *bite* is by no means clear or stable over time, and possibly depends on the "political mood"); and (iii) a necessary condition for identification of the switch in a regime-switching model is the observation of independent variation in the exogenous variables, such as the marginal cost of imports and the marginal cost of the domestic industry. When one moves beyond the present imports threat to consider other types of constraints (recall the Introduction), finding instruments of this kind may in practice be too difficult.

constrained, the coefficient on $-q\frac{\partial p(q)}{\partial q}$ will be biased toward zero. In general, unless the researcher is somehow sure that *none* of his observations correspond to constrained outcomes, his estimates will potentially be inconsistent. An example is provided by Genesove and Mullin's (1998) seminal test of the standard methodology using data from the sugar industry. They conclude that the methodology "performs reasonably well in estimating θ" (p. 370), but they do obtain a bias, $\hat{\theta} < \theta$, albeit small in their context. Proposition 1 and its corollary suggest that to the extent that market outcomes in the sugar industry were constrained by the threat of entry, this might lead to a downward bias in the estimated degree of market power[16].

3   Industry and data[17]

Brazil ranks sixth among cement-producing countries, with output of approximately 40 mtpa (million tonnes p.a.) in the period 1998 to 2000. As shown in Figure 4, in 1999 57 active plants were scattered across a geographic area slightly smaller than that of the US[18]. This spatial distribution is not even, however, as consumer markets and thus plants are heavily concentrated in states located along (or in proximity to) an extensive Atlantic coastline, in particular the relatively wealthy and populated states in the Southeast and South regions of the country[19]. States to the northwest of the center of the country are sparsely populated and are largely covered with jungle. The industry also exhibits high concentration in terms of ownership. Though the 57 plants in 1999 were owned by 12 firms in total[20], the largest firm Votorantim commanded a nationwide shipment share of 41%, followed by Grupo João Santos (GJS), Holcim, and Lafarge, with respective shipment shares of 12%, 9% and 8%. As Figure 5 indicates, this national picture hides a lot of variation at the local, statewide level.

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[16]Genesove and Mullin (1998) acknowledge the price-restraining effect of potential imports: "Although very little refined sugar was ever imported into the United States, in the early years of the Sugar Trust the threat of European imports affected U.S. prices. In 1888 and 1894, Havemeyer acknowledged setting the price of refined sugar so that none would be imported from Europe" (p. 358; the Sugar Trust was the largest firm, with a 63% market share, and Havemeyer was its president). Given such high concentration, the direct measure of market power θ (which stems from a low observed price-cost mark-up $\frac{p-c}{p}$, and a moderate elasticity η of around -1.05 for most part of the year) is surprisingly low, the authors attributing this to the threat of entry.

[17]This section covers some key features of the Brazilian cement industry and of the data. Further considerations of the industry, as well as a detailed account of the sources and treatment of the data, are provided in a separate *Supplementary Section*.

[18]With a population corresponding to two-thirds that of the US, cement consumption per capita in Brazil amounts to 232 kg as compared to 415 kg in the US (SNIC 2002).

[19]The Federative Republic of Brazil is a federation of 27 states. The coastal states are those running clockwise from the north-most point of the country – the state of Amapá (AP) – to the south-most state of Rio Grande do Sul (RS). The long Atlantic coastline measures 7,491 km (4,655 miles).

[20]The 1990s saw growing consolidation of assets in the industry, particularly during 1995 to 1997, years of sharp growth in cement consumption, as I note below. Compared to the 12 firms that ran operations by 1999, the industry had consisted of 19 producers in 1991.

The Brazilian cement industry experienced two distinct periods over the 1990s. Up until mid 1994, a period of very high inflation and low macroeconomic growth, cement consumption was stagnant at around 25 mtpa. With the successful implementation of the *Real* economic stabilization plan in July 1994, cement consumption resumed its growth at a rate of 10% p.a., reaching 40 mtpa by 1998-99, pulled by exogenous growth in the construction sector[21]. Thanks to the industry's practice of investing in idle capacity, firms were able to meet rising demand[22].

Given the short shelf life of cement, firms produce for immediate consumption. Shipments from producer plants to buyers in local consumer markets is primarily (i.e. 90%) carried out by road – as opposed to rail or water. In line with other developing countries, and in contrast to developed nations, around 80% of volume is shipped in bags to resellers who then sell on to small-scale consumers; only 20% is shipped in bulk by the industry directly to consumers, usually ready-mix concrete firms, large construction firms or producers of concrete products.

The role of imports in Brazil   Imported cement (including the intermediate product clinker[23]) constitutes a small share of domestic consumption. As shown in Figure 6, in the period 1989 to 2003, this share has amounted to at most 2-3% of consumption across Brazil. This low level stands in striking contrast to the penetration of imports in the US. Carlsson (2001) reports that "imports represent a substantial and increasing part of the market in the United States, ranging between 10 and 17 percent of domestic consumption since 1985" (p. 7). The share of imports in some coastal US markets is actually as high as 30% (exported from regions as far away as Southeast Asia)[24].

Despite paling in comparison to the US, the limited penetration of imports into Brazil's (essentially coastal) markets hides their role in restraining domestic prices. The trade liberalizing reforms of the early 1990s opened the door to the threat posed by the

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[21] The sharp slowdown in inflation brought about by the *Real* stabilization plan represented a reduction in the transfers from the private sector to the government (in the form of an "inflation tax"). In particular, the large mass of low-income households who previously had no access to instruments of monetary protection, such as price-indexed savings accounts, saw a significant rise in real incomes. Given their high propensity to consume, this boosted the demand for consumer goods – notably food, clothing and durables – and the demand for housing.

[22] Capacity utilization typically hovers around 65%. As I present shortly, the domestic industry was able to keep imports at bay even during the construction boom of 1995 to 1997. Also during this period, a strong appreciation of the local currency was accompanied by a fall in domestic cement prices. Capacity is added mostly to existing plants: over the 1990s, only four new (fully-integrated, i.e. comprising limestone quarry, kiln and grinding mill) plants were set up, invariably by incumbents.

[23] Clinker is the main component of ordinary cement.

[24] Despite its bulkiness relative to price, the development of specialized seaborne handling and transportation equipment from the 1970s has led to cement being traded internationally between regions with access to sea (see Dumez and Jeunemaître 2000 for a historical account). As for cement exports from both the US and Brazil, these account for less than 1% of domestic production (though in Brazil the current trend is upwards).

entry of imports. Two stylized facts concerning prices are consistent with the imports-constraining story. First, domestic cement prices in local currency (the *real*, R\$) are highly correlated with the price of the US dollar in local currency (i.e. the exchange rate). To provide a flavor, Figure 7 depicts the evolution of cement prices in the state of Rio Grande do Sul – where one-firm and two-firm concentration ratios respectively amounted to 55% and 84% in 1999 – both in (current) local currency and in US dollars. Despite the occurrence of large swings in the exchange rate, the domestic cement price converted into US dollars is quite steady since 1995[25] [26]. Second, in the cross-section of local markets, cement prices are increasing in the market's distance from the coast. However, I do not take either of these two stylized facts as *prima facie* evidence that potential imports restrain prices. The reason is that these stylized facts could also be explained by alternative stories, such as factor prices (i.e. oil) being set in hard currency on the world market, and producers incurring higher transport costs to distribute cement in less densely populated areas. The estimation of a very low market price elasticity of demand in equilibrium, coupled with high price-cost margins, and backed up by interview evidence, will be the key elements in support of my claim.

Data available: Plant-to-market cement flows and delivered marginal cost
On the demand side, I observe monthly cement consumption and consumer prices (i.e. retail prices set by resellers) for each of the 27 states in the period 1991 to 2003. As proxies for the exogenous demand for cement, I use alternative series of economic activity, either in the construction and building sector or aggregated across sectors of the economy[27].

The truly distinctive feature of the data is the supply side. Most unusually, I observe

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[25]Until January 1999, Brazil had a fixed (and overvalued) exchange rate regime. The currency was then floated, in the midst of the "Brazil currency crisis", almost instantly depreciating by 70% against the US dollar, but later partially receding. Other periods of exchange-rate instability took place in 2001 (commonly attributed to the Argentina crisis next door) and in the second half of 2002, with the uncertainty surrounding the presidential election late that year. The relatively flat evolution of domestic cement prices in US dollars is consistent with imports setting a price ceiling at around 6-7 US dollars per bag (this would correspond to the US-dollar equivalent of c^I in Section 2). That it seems to take domestic producers between 6-12 months to raise domestic prices back to this ceiling in US dollars upon large unexpected bouts of devaluation suggests that raising domestic prices in local currency is not friction free (perhaps the industry is wary of attracting negative publicity).

[26]To illustrate, equity analyst Zaghen (1997) writes: "(a)lthough imports accounted for only 1.6% of the Brazilian total consumption in 1995, reaching 451.3 thousand tons, it represents a constant threat to domestic producers, pressing down domestic prices and imposing a price ceiling of US\$ 70 per ton" (p. 24). (The author refers to the landed price of imports at the port of entry.) Further evidence that domestic producers were threatened by imports is their successful lobbying of government in passing antidumping measures – namely a 23% import tariff – in the late 1990s against Venezuelan and Mexican cement producers who were starting to make inroads into local markets particularly in the north and northeast of the country.

[27]While cement is a critical input to construction, it accounts for only a small share of construction budgets. Taking construction activity to exogenously move the demand curve for cement (or, equivalently, concrete) is typical in the literature (e.g. Syverson 2004 uses construction sector employment).

each plant's shipments to each and every state, enabling me to map the flow of cement from the plant to the consumer. I thus observe the level at which each plant, and thus firm (given observed plant ownership), supplies (or does not supply) each local market (state). To calculate delivered marginal cost for every plant-market pair over time, I combine plant-to-market shipments with observed (i) plant-level characteristics (e.g. location, capacity, number and age of kilns, type of fuel usage, and the proportion of shipments in bags as opposed to bulk); (ii) engineering estimates (in view of the simple fixed-coefficients technology of cement production and distribution); and (iii) local factor prices (e.g. fuel oil, coal, electricity, wages, and freight rates).

A few comments are in order. While I do not directly observe freight rates paid by cement producers, I proxy for these using data on freight rates for agricultural commodities collected over the period 1997 to 2003 for thousands of different routes across Brazil[28]. Similarly, I do not observe producer prices for cement, but I can back these out from consumer prices based on the high proportional sales taxes and assuming retailers (i.e. resellers) are competitive[29]. Delivered marginal costs thus include the entire supply chain from the producer of cement to the retail consumer, encompassing the reseller: in addition to plant marginal cost, total plant-to-market marginal cost consists of plant-to-market freight, sales taxes and the reseller's mark-up[30].

Figure 8 depicts cement prices (in R\$ for the standard 50 kg bag, at a constant December 1999 level[31]), cement consumption and construction activity from January 1991 to December 2003 for the largest market, the state of São Paulo (*SP*). The month in which the stabilization plan was enacted, July 1994 (month 43), is marked in each plot with a dotted line. Following the lifting of price controls in November 1991, prices approximately doubled, remaining in the high R\$ 14 to R\$ 16 range until 1994. In the post-stabilization period they gradually declined back to R\$ 7 by late 1996, gradually rising thereafter. The sharp increase in consumption within two years of stabilization, from a monthly level of 600 mt to 1000 mt, pulled by a 20% jump in construction activity, is evident from the plots. Some factor prices are also portrayed. Of note, there is high correlation in the post-stabilization phase between the price of cement and the prices of fuel oil and diesel oil[32]. This is expected in view of (i) my earlier claim (at this point) that imports set a price ceiling for cement and thus the price of cement (in local

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[28] The transportation of, say, soyabean and maize are close substitutes in the supply of cement freight (Soares and Caixeta Filho 1996). To emphasize, see the *Supplementary Section* for further comments.

[29] The assumption of competition among resellers follows from several field interviews of salespeople at producers and buyers at resellers. I also verify this assumption by, for example, comparing the backed-out producer prices to producer prices that I was able to obtain directly from some producers.

[30] I also argue that my calculations of marginal cost may actually overstate the true marginal costs. But when I turn to the testing of conduct, such a bias would only reinforce the results.

[31] I use an economy-wide General Price Index (GPI). Owing to the high levels of inflation prevailing in the first 42 months (out of 156) of the sample, particular attention has been paid to the conversion of current prices to constant prices. (Figure 7, in contrast, presents current prices.)

[32] Fuel oil and diesel oil, used respectively in production and in transportation, are the two major

currency) is highly correlated with the exchange rate, and (ii) oil is a global commodity and policy in the oil sector from the second half of the 1990s has prescribed domestic oil prices varying in line with the world price (and hence with the exchange rate).

A glance at price-cost margins   With respect to firm profitability, Figure 9 shows the evolution of average consumer prices, marginal cost and price-cost margins on the leading firm Votorantim's actual sales across Brazil, in constant R$ per bag. Prices and marginal cost have been increasing since late 1996, the latter owing chiefly to increases in the price of fuel oil and diesel (freight) and the fact that sales taxes are proportional to prices – recall that cost relates to the entire supply chain. The picture is similar across firms. *In sum, the industry wields considerable market power, despite the threat of import competition.* Across producers, across states and over time, the price-cost margin as a proportion of the consumer price lies in the region of 25-45% (equivalent to 40-65% as a proportion of the producer price net of sales tax)[33].

4   Inferring demand and conduct in the Brazilian cement industry

4.1   A "road map"

Section 4.2 estimates demand in each local market (state), obtaining very low market price elasticities of demand, of the order of -0.5. Such low market elasticities are systematic across local markets, including markets where the one-firm concentration ratio is as high as 80%. Two main possibilities arise to rationalize why an industry facing such inelastic demand does not cut output to raise prices to a point where demand is more elastic[34]: (i) strategic behavior among incumbent firms is such that there is weak pricing power[35], or (ii) prices are constrained by the potential behavior of agents other than

components of cost. From July 1994, correlation coefficients (all highly significant) are as follows: 0.72 between the price of cement and the (US dollar) exchange rate; 0.86 between the price of cement and the price of fuel oil; 0.77 between the price of fuel oil and the exchange rate.

[33]I conduct two robustness tests of the calculated marginal costs and the resulting price-cost margins. The first one is based on unusually-detailed accounting data reported by country of operation (and by line of business) by the multinational firm Cimpor. The second test is based on accounting data sampled among establishments in the cement industry by the Brazilian Institute for Geography and Statistics (IBGE) as part of their Annual Industry Survey (PIA) series.

[34]A third possibility hinges on a very special class of models of spatial competition, à la Hotelling-Salop, where a firm sets only a "mill" price and cannot price discriminate over space. The restrictive nature of pricing then ensures that a low *market* price elasticity of demand does not translate into a low price elasticity of demand faced by the *firm*. See Salvo (2005).

[35]For example, behavior is competitive, or low concentration ensures 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 (thus in equilibrium the price remains at a level consistent with aggregate demand being inelastic).

consumers (market demand) and incumbents, such as entrants or regulators. By the first alternative, an industry seeing such inelastic demand would not be able to restrict output to raise prices because competition among incumbent producers drives prices down toward marginal cost. However, I reject competition on the basis of the large observed price-cost margins, amounting to around 50% of producer prices. The second explanation is the accepted hypothesis. While market demand in equilibrium is inelastic, the residual demand which the domestic industry faces at the price ceiling posed by high-cost imports is highly elastic[36]. Attempts by the domestic industry, already enjoying a large price-cost margin, to raise prices above this ceiling would only invite foreign entry[37]. This imports-arbitrage hypothesis is further supported by a wealth of anecdotal and interview-based evidence. It is also consistent, as argued in Section 3, with the high correlation between cement prices and the exchange rate.

In this constrained setting, Section 4.3 then illustrates the poor performance of the standard methodology for inferring supply. Assuming costs are unknown, I overlook the latent effect of imports and impose the regular moment conditions for identification of the standard (misspecified) pricing equation. The conduct parameter is estimated to be close to zero and costs are estimated to be close to prices, wrongly suggesting that the outcomes in the Brazilian cement industry are competitive. Estimates of the true price-cost margins are severely biased downward. Intuitively, the underlying competitive constraint posed by the fringe of *high-cost* foreign entrants on any domestic firm's supply decision is being misinterpreted as a competitive constraint posed by the firm's *low-cost* domestic rivals.

Finally, Section 4.4 uses observed marginal cost to show that firm-level behavior is considerably more collusive as benchmarked against a Cournot model. I find that Brazilian cement firms are tacitly dividing geographic markets. In view of the aggregate imports constraint, the test I design uses firm-level supply data to each local market to place a tighter lower bound on the collusiveness of firm conduct than would be possible with market-level data alone.

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[36]That market demand is inelastic in equilibrium owes to demand, costs and firm conduct (i.e. the structural parameters of the data generating process) being such that this upper limit to prices binds, and demand happens to be inelastic at this limit, as in the constrained monopoly of Figure 2.

[37]Other studies of cement have found low market price elasticities. For example, Röller and Steen (2006) find an elasticity of -0.46 for Norway, while Jans and Rosenbaum (1996) report an average -0.81 across 25 regional US markets. It is conceivable that in these markets imports also restrain domestic oligopolies. The explanation commonly advanced behind such inelastic demand is that cement accounts for a low share of construction budgets and that it has few substitutes (except in highway construction, where asphalt is a substitute). Yet while helping to explain the steepness of the inverse demand curve, this does not explain the steepness *at the equilibrium*. One must still explain why an industry, facing such inelastic demand at the market price, does not cut output in an attempt to raise prices and thus move up along the demand curve to a point where demand is more elastic.

4.2   Demand

There are L (geographic) markets (identified with states of the Brazilian federation), indexed by $l = 1, ..., L$. Scattered across these L markets are I plants, indexed by $i = 1, ..., I$. Let $i = 0$ index the aggregate fringe of foreign suppliers. The flow of cement is contained in a set of $(I + 1) \times L$ 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. Let q_{lt} denote total shipments to market l in period t, i.e. consumption; then $q_{lt} = \sum_{i=0}^{I} q_{ilt}$. The demand function (1) in each market l is:

$$q_{lt} = D(p_{lt}, Y_{lt}, \varepsilon_{lt}^d; \alpha_l) \tag{13}$$

where p_{lt} is the price of cement to the consumer, Y_{lt} are exogenous variables moving demand (i.e. construction activity), α_l are market-specific parameters to be estimated and ε_{lt}^d is the econometric error term.

Estimation of (13) must deal with the (potential) endogeneity of prices. The choice of instruments will depend on whether the imports constraint binds, which in turn depends on the behavior of domestic firms. There are therefore two situations to consider.

Identification 1: Imports-arbitrage constraint binds at the industry equilibrium   In practice, due to the presence of frictions, cement prices will not be exactly equal to the marginal cost of imports c^I. Prices and c^I should be highly correlated however. As mentioned in Section 2 (recall the right panel of Figure 2), fluctuations in c^I allow one to trace out the demand curve (assuming c^I does not rise to the extent where imports no longer have bite). c^I is a function of factors such as the exchange rate, world fuel prices (used in the production of clinker abroad and in the international transport of cement), tariffs and port handling charges, and domestic freight to the consumer (the latter being highly correlated with the domestic price of diesel oil). Observed factors such as the exchange rate, world oil prices and domestic diesel oil prices (all in local currency in constant terms) can then instrument for prices in the estimation of (13) (under the identifying assumption that these factors are not correlated with the unobserved market-specific demand shocks ε_{lt}^d).

To the extent that the "frictions component" of cement prices – i.e. the part of p not determined by c^I – is orthogonal to ε_{lt}^d, prices can be treated as predetermined and (13) can be estimated by OLS[38].

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[38] The model I have in mind when imports restrain domestic prices is as follows. Cement prices p are determined by the marginal cost of imports c^I and a frictions component ζ, i.e. $p = c^I + \zeta$. As for c^I, as the econometrician I observe some cost drivers V_1 but not others V_2, where $c^I = V_1 \kappa + V_2 \phi$, and κ and ϕ are parameters. Under the identifying assumption that $E(V_1 \varepsilon^d) = 0$, V_1 (e.g. the exchange rate)

Identification 2: Imports-arbitrage constraint does *not* bind at the industry equilibrium When imports do not restrain domestic prices, traditional cost-shifters may be used to instrument for cement prices. These include factor prices (i.e. prices of kiln fuel such as fuel oil and coal, electricity prices which determine the cost of grinding, the price of diesel oil which drives the cost of freight, and wages, the latter also impacting freight in addition to the cost of production) and other supply-shifters such as plant capacity, to the extent that changes to scale impact marginal cost.

Demand specification I fit alternative parametric specifications for the market-level demand function (13), such as the loglinear form[39]:

$$\log q_{lt} = \alpha_{1l} + \alpha_{2l} Y_{lt} + \alpha_{3l} \log p_{lt} + \alpha_{4l} Y_{lt} \log p_{lt} + \varepsilon_{lt}^d \tag{14}$$

The inclusion of an interaction term between (log) price and the exogenous demand variable, $Y_{lt} \log p_{lt}$, allows the demand curve in logs to rotate – in addition to shift, through the level term Y_{lt} – as exogenous demand varies. By the above discussion, (14) is estimated, for each local market, by (I) OLS, (II) 2SLS using the exchange rate and other prices relevant to the marginal cost of imports as instruments, and (III) 2SLS using factor prices as instruments. These three sets of results are depicted in Figure 10 for the state of São Paulo (SP), denoted respectively as "OLS", "IV imports bite" and "IV imports no bite". Most estimated coefficients are significantly different from zero, many at the 1% level of significance. The interaction term is found to be negative and highly significant: the demand curve (in logs) rotates anticlockwise as exogenous demand expands. Taking the mean values for exogenous demand in the two respective periods ($\bar{Y}_{SP,pre} = 2883$ and $\bar{Y}_{SP,post} = 3338$), the market price elasticity of demand during the pre-stabilization phase amounts to (an inelastic) -0.17, rising to -0.33 during the post-stabilization phase. Thus, as prices in the economy stabilize and an average 16% exogenous increase in the demand for cement occurs, the price elasticity doubles from around -0.2 to around -0.4. Figure 11 plots the fitted demand curve

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can be used to instrument for prices in the estimation of (13). In addition, if $E(\zeta \varepsilon^d) = 0$ (and of course $E(V_2 \varepsilon^d) = 0$ as well), equation (13) can be estimated consistently by OLS.

[39]Some comments on the results that follow are in order (see the *Supplementary Section* for further details): (i) the loglinear specification estimates that I present are robust to the choice of alternative functional forms; (ii) I provide detailed estimates for the largest market, the state of SP, but later show that results follow a common pattern across states; (iii) for each market there are 156 monthly observations, from January 1991 to December 2003; (iv) three quarterly dummies are included to capture the seasonality of sales; (v) the finding of low elasticities is robust to estimating using the post-stabilization subsample only (the latter 114 observations); and (vi) standard errors are heteroskedasticity and autocorrelation-robust (1-lag Newey-West errors). The *Supplementary Section* discusses specification tests such as heteroskedasticity and serial correlation, and overidentifying restrictions, as well as further robustness tests (e.g. using fixed-effects IV panel data estimation, reversing the dependent variable). To check whether overidentification may be driving efficiency at the expense of consistency, I reestimate baseline regressions using subsets of the sets of instruments, to obtain similar elasticities – see regressions (II B) and (III B) in Figure 10.

evaluated at the means for exogenous demand in the pre- and post-stabilization phases. In addition to the state of São Paulo (SP), plots are drawn for the three next largest markets, the states of Minas Gerais (MG), Rio de Janeiro (RJ) and Bahia (BA). Again, as stabilization takes place and exogenous demand grows, the demand curves shift out and rotate anticlockwise. This suggests that there may be a typical pattern across states.

Results by state Figure 12 summarizes results across 17 states, from regression (II), again using the full sample (pre- and post-stabilization phases). (I drop 10 sparsely populated northwestern states from the analysis owing to measurement error[40].) The pattern is similar and consistent with the results reported for the state of SP. Evaluating exogenous demand at its mean value in the post-stabilization phase, the price elasticity of demand is negative for all 17 states, and significant at the 1% level in 15 states. Post-stabilization price elasticities vary from a minimum (in absolute) of -0.14 to a maximum of -0.72, with a mean of -0.41 and a standard deviation of 0.14. Elasticities are low even in states where the supply of cement is highly concentrated, such as the state of Santa Catarina (SC), in which the one-firm concentration ratio is 78% (in 1999). The average price elasticity in the pre-stabilization phase is negative in 16 out of 17 states, 9 of which are significant at the 10% level or higher. The mean pre-stabilization price elasticity across states is lower: -0.22.

In sum, regardless of the type of price instruments employed (or using prices themselves, under OLS), the choice of which depends on whether market outcomes are constrained by imports, I estimate very low market price elasticities of demand, of the order of -0.5.

4.3 The standard methodology: Inconsistent supply estimates

I proceed to estimating a pricing equation such as (2), as is standard in the empirical literature on market power[41]. In order to complete the specification of the structural econometric model, define plant i's costs as

$$C_{it} = C(q_{it}, \mathbf{q}_{it}, W_{it}, V_{it}, \varepsilon_{it}^s; \beta)$$

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[40] As a point of comparison, the global market research firm ACNielsen, that has long been established in Brazil, does not audit these jungle states owing to their unusual geo-demographic characteristics. The 10 states together account for 60% of Brazil's land mass but only 11% of its cement consumption. Note, however, that when I later analyze supply decisions to the 17 remaining consumer markets, I do consider sourcing from plants located in *all* 27 states.

[41] I now ignore that the true model is (4) and that the estimated model is thus (5) with $E(Y'\xi^s) < 0$. Since I take the view of a researcher who overlooks the price-constraining effect of imports, thinking that he is specifying (2) with $E(Y'\varepsilon^s) = 0$, I refer to the pricing equation error as ε^s and not ξ^s.

where $q_{it} := \sum_{l=1}^{L} q_{ilt}$ denotes plant i's shipments aggregated across markets l (equal to production), W_{it} are the prices it pays for its factors, and V_{it} are other exogenous variables that shift supply. Note that costs by plant will not only depend on the plant's total shipments q_{it} but also on the destination of these shipments $\mathbf{q}_{it} := (q_{i1t}, q_{i2t}, ..., q_{iLt})$, owing to market-specific factors such as freight. β is a vector of common parameters to be estimated and ε_{it}^s is a plant-specific error. The I plants are owned by F firms, indexed by $f = 1, ..., F$. Define $\mathcal{O}_{ft}$ as the set of plants owned by firm f in month t. Writing the market-level demand function (13) in inverse form $p_{lt} = p(q_{lt}, .)$, in period t firm f solves

$$\max_{q_{ilt} | i \in \mathcal{O}_{ft}, \forall l} \sum_{l=1}^{L} \left[p(q_{lt}, .) \left(\sum_{i \in \mathcal{O}_{ft}} q_{ilt} \right) \right] - \sum_{i \in \mathcal{O}_{ft}} C(q_{it}, \mathbf{q}_{it}, .)$$

In words, firm f sets shipments from each plant it owns to each market to maximize its profits, which correspond to the difference between the sum of revenues across markets and the sum of costs across plants. Denote the derivatives of the (inverse) demand and cost functions with respect to q_{lt} and q_{ilt} respectively as $p_1(.)$ and $c(.)$. Following Bresnahan (1989), the first-order condition for multi-plant firm f with regard to shipments from its plant $i \in \mathcal{O}_{ft}$ to market l, i.e. q_{ilt}, yields a pricing equation for each plant i - market l pair:

$$p_{lt} + p_1(q_{lt}, Y_{lt}, \varepsilon_{lt}^d; \alpha_l) q_{lt} \theta_{flt} \leq c(q_{it}, \mathbf{q}_{it}, W_{it}, V_{it}, \varepsilon_{it}^s; \beta)$$

where θ_{flt} is a firm-level conduct parameter. For $q_{ilt} > 0$ (i.e. an interior solution) and specifying an additive econometric pricing error, one may implement this pricing equation as

$$p_{lt} = -\theta_{flt} \frac{p_{lt}}{\eta_{lt}} + c(q_{it}, \mathbf{q}_{it}, W_{it}, V_{it}; \beta) + \varepsilon_{ilt}^s \tag{15}$$

In what follows, I present estimation results corresponding to a market-level counterpart to the plant-level pricing equation (15)[42]. In view of the fixed-coefficient technology of production, I specify average market marginal cost c as being linear in average market factor prices W_{lt} (namely fuel oil, coal, electricity, labor and freight[43]) and flat in

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[42] As mentioned in footnote 11, the market-level equation should be viewed as an average across plants' pricing equations. Owing to the lack of firm-level data, most empirical IO studies have no choice but to estimate a market-level equation. Though I have the luxury of observing plant-level data, I here choose to follow suit to simplify the exposition. Importantly, I have estimated a plant-level pricing equation and have ensured robustness of the conclusions I derive from what follows.

[43] The average factor price, say that of electricity, for a given market is calculated as the average price of that factor paid by the plants sourcing that market (weighted by the sourcing plants' shipments to that market). The price effect of substitute kiln fuels (fuel oil and coal) are interactions of the average price of the fuel and the average use of that fuel in the production of cement shipped to the market (i.e. given the location of coal mines in the south of the country, coal prices have a larger effect on the cost of cement plants located in the south). A market's average plant-to-market freight price is modeled as the

quantity (though in other specifications I have also allowed average market marginal cost to vary in quantity). I allow cost to shift according to the (shipment-weighted) average size and age of the plants shipping into the market, V_{lt} (e.g. marginal cost in a market served by high-capacity plants should be lower). (Finally, a dummy is included to account for price controls in the first ten months of 1991: this may be viewed as an additional supply-shifter V_{lt}.) The market-level pricing equation is thus

$$p_{lt} = -\theta_l \frac{p_{lt}}{\hat{\eta}_{lt}} + W_{lt}\beta_1 + V_{lt}\beta_2 + \nu_l + \varepsilon_{lt}^s \tag{16}$$

where ν_l is a market-specific fixed effect and the market-specific conduct parameter θ_l is time-invariant (other specifications I have fitted allow θ_l to vary over time, such as upon stabilization). Equation (16) is fitted using fixed-effects instrumental variables panel data estimation, where the endogenous regressor $\frac{p_{lt}}{\hat{\eta}_{lt}}$ is instrumented using excluded exogenous demand variables Y_{lt}, and thus the orthogonality condition $E(Y'\varepsilon^s) = 0$ is imposed. Since the elasticity $\hat{\eta}_{lt}$ is estimated (Section 4.2) rather than known, I compute bootstrapped (heteroskedasticity-robust) standard errors (with 1000 repetitions, reestimating demand in the first stage for each bootstrap sample) for the fitted coefficients $\hat{\theta}_l$ and $\hat{\beta}$ (and $\hat{\nu}_l$). Notice that though knowledge of the nature of technology is used when specifying marginal cost to be linear in factor prices, *marginal cost is assumed to be unknown*: this is estimated from the observed supply-shifters (W_{lt}, V_{lt}) and the estimates of the fixed coefficients (β, ν) as $W_{lt}\hat{\beta}_1 + V_{lt}\hat{\beta}_2 + \hat{\nu}_l$.

Figure 13 reports estimation results[44]. The coefficients on the prices of fuel oil, coal, electricity and freight are all positive and significant. The coefficient on the average size (resp. age) of plants is negative (resp. positive), as expected, though not significant. On the other hand, contrary to intuition, the price of labor is significantly negative[45]. The price-cost margins are estimated to be very low; these are pictured in Figure 14, along with 95% confidence intervals, for the state of Rio de Janeiro (RJ), for example. The dual to these cost estimates are the low estimated conduct parameters $\hat{\theta}_l$, not significantly different from 0, suggesting competition[46]. For the state of RJ, a $\hat{\theta}$ of 0.0079 would correspond to the equilibrium price-cost margins of a static symmetric 130-firm Cournot industry (1/0.0079). Dividing $\hat{\theta}_{RJ}$ by the (negative of the) estimated elasticity $\hat{\eta}_{RJ}$ of -0.48 from Figure 12, the estimated (average) price-cost margin as a proportion of price is only $0.0079/0.48 \approx 1.6\%$ (recall expression (3)).

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interaction of the average distance to market across the plants sourcing that market (again weighted by shipments) and a transport price index (based heavily on the price of diesel oil).

[44]The comments that follow refer to estimation (I), based on the entire period, though estimates based only on observations from the post-stabilization period are provided to demonstrate robustness of the conclusion that follows.

[45]One can, however, attempt to rationalize this result through a rent-sharing story (e.g. Clark 1980).

[46]While the estimated confidence intervals for $\hat{\theta}_l$ vary according to the specification (such as the functional form for demand), low (absolute) values are a robust result.

It is clear from our knowledge of marginal cost and price-cost margins in the industry that these estimates are inconsistent. Figure 14 also depicts the (much higher) direct measures of (average) price-cost margins on sales to the state of *RJ*. What lies behind the market price elasticities of demand of the order of only -0.5 in equilibrium, *is not the prevalence of actual competition, as suggested by* $\hat{\theta}$, *but the constraining effect of potential high-cost competition.* The standard identifying assumption's failure to hold results in the underestimation of the degree of market power (Corollary 1). (Indeed, the p-values of overidentification tests à la Sargan and Hansen – where the null is that the set of instruments is valid – is 0.0000 for any overidentifying set.) The finding that the coefficients on factor prices and other supply-shifters are of the expected sign (bar wages) and mostly significant may lead one to misjudge that the econometric model is appropriately specified. But the estimated coefficients are only picking up the expected correlation between cement prices and factor prices. They are not consistent estimates of the structural cost parameters β.[47]

4.4 Inferring market division in a constrained equilibrium when marginal cost is observed

How can one use observed marginal cost to identify firm-level behavior underlying a constrained market equilibrium? Clearly, a direct comparison of marginal cost to price provides a test of perfectly competitive behavior against less competitive models of conduct. But, other than perfect competition, how may one distinguish empirically between alternative less competitive models of firm behavior when the threat of entry constrains these alternative models to generate the same *aggregate* outcome? I design a test that uses firm-level supply data to each local market to place a tighter bound on firm conduct than would be possible from market-level data. It adopts the Cournot behavioral model as a benchmark. While market outcomes may be consistent with Cournot, in that the price ceiling $p = c^I$ (equivalently, output floor $q = p^{-1}(c^I)$) posed by the imports constraint binds at the market-level Cournot solution, I may still be able to reject Cournot at the firm level, in favor of "more collusive" conduct, for a

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[47] A comment on this particular industry where conditions such as a steep demand curve result in demand being inelastic at the equilibrium. Assume one overlooks the binding imports constraint and thus misguidedly considers the class of behavioral models nested in the static pricing equation $p + \theta_f \frac{p}{\eta} = c_f$. Clearly, an η of -0.5 is not consistent with cartel behavior ($\theta = 1$): a cartel would cut output until its marginal revenue were equal to marginal cost (and thus positive). Nor will such a low value of η be consistent with Cournot, unless all firms have small market shares (i.e. if the largest firm has a 50% market share, say, then under Cournot $\max_f\{\theta_f\} = 0.5$ implying that this firm's marginal revenue is zero). But concentration in Brazilian cement is anything but low. Any statistical model selection exercise à la Gasmi, Laffont and Vuong (1990, 1992) from this misspecified set of alternative models will thus result in, say, both the cartel and Cournot models being rejected in favor of price-taking behavior and zero price-cost margins ($\theta = 0$). One's judgement that this price-taking hypothesis is appropriate would only be reinforced by the good fit of an OLS regression of cement prices on factor prices (and other supply-shifters, along with a set of market dummies): R^2 is 54%!

considerable proportion of firms' supply decisions, should the true behavioral model be one where firms tacitly collude to divide geographic markets[48]. In other words, the test will have greater power to the extent that, say, firm 1 tacitly agrees to give firm 2 the upper hand in market B in exchange for the latter staying away from market A.

4.4.1 Testing Cournot against "more collusive" conduct

Figure 15 establishes the Cournot benchmark in any given local market[49]. In the left panel, the steep line depicts firm f's reaction function in the absence of imports; write this as $q_f = R_f(q_{-f})$, where $q_{-f} := \sum_{j \neq f} q_j$ is the joint output of firm f's (domestic) rivals. The line with slope -1 represents the imports constraint, given by $q_f + q_{-f} \geq p^{-1}(c^I)$. Thus, under the null, firm f's "constrained" best response is the outer envelope

$$q_f = \tilde{R}_f(q_{-f}; c^I) := \max(R_f(q_{-f}), p^{-1}(c^I) - q_{-f})$$

The constrained reaction function consists of two segments. For high enough q_{-f} such that $\tilde{R}_f(q_{-f}; c^I) + q_{-f} \geq p^{-1}(c^I)$, imports have no bite and the standard Cournot pricing equation holds. Now, for lower q_{-f}, the imports constraint binds and firm f's optimal reply *exceeds* the quantity that it would set in the absence of imports. The following proposition thus holds:

Proposition 2 *("Constrained" Cournot first-order condition and sufficient statistic to reject Cournot behavior) In the presence of imports, if firm f behaves as a Cournot player, it will be the case that*

$$p(q) + \frac{p(q)}{\eta(q)} \frac{q_f}{q} \leq c_f \tag{17}$$

This condition holds as a strict inequality when the imports-arbitrage constraint $q_f + q_{-f} \geq p^{-1}(c^I)$ binds, in which case price is equal to the marginal cost of imports c^I.[50] This translates into the following test. Rewrite (17) as an equality:

$$p(q) + \frac{p(q)}{\eta(q)} \frac{q_f}{q} = \varphi_f + c_f \tag{18}$$

Under the null of Cournot behavior, $\varphi_f \leq 0$. When the imports constraint binds, $\varphi_f < 0$ is consistent with Cournot behavior. The finding that $\varphi_f > 0$ allows one to reject the

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[48] "More collusive" behavior is employed in the sense that, were the constraint not to bind, aggregate output supported by such behavior would be lower than the Cournot oligopoly solution.

[49] For a model with a similar flavor where a Cournot oligopoly may deter entry by producing the limit output, see Gilbert and Vives (1986). As in the empirical literature on conduct, the Cournot assumption serves as a benchmark.

[50] Condition (17) also holds as an inequality in the case of a corner solution (i.e. $p(q_{-f}) < c_f$ such that $q_f = R_f(q_{-f}) = 0$), but since this case is standard I omit it from the proposition.

hypothesis that firm f is behaving in Cournot fashion, in favor of more collusive behavior, regardless of whether the imports constraint binds or not.

The intuition for why $\varphi_f \leq 0$ does not allow us to reject Cournot in favor of *less* collusive (or "more competitive") behavior is that cutting output when the imports constraint binds would only open the door to imports. This is illustrated in the right panel of Figure 15, which also depicts the rival firms' constrained joint reaction function, $q_{-f} = \tilde{R}_{-f}(q_f; c^I)$, now relabeled as (duopolist) firm g. Say that in a spatial setting, while the imports constraint binds, firms are dividing markets and the observed outcome in a particular local market, marked "•", is one where firm-level shipments to that market are "sufficiently" asymmetric[51]. While firm g's behavior (toward the local market) is still consistent with Cournot (since under the null it does not cut output since the imports constraint is binding), one can reject the hypothesis that firm f's is behaving in Cournot fashion in favor of more collusive behavior. Firm f is restricting output as compared to the supply decision of a constrained Cournot firm. The point is to recognize that for a Cournot firm, the general (i.e. allowing potential imports arbitrage) pricing condition (17) has to hold. That is, for no Cournot firm can (perceived) marginal revenue *exceed* marginal cost, otherwise the firm would optimally expand supply, and this holds irrespective of whether the imports constraint binds or not. Notice that the set of constrained collusive equilibrium outcomes is collinear with the boundary to the imports constraint and is a superset of the set of constrained Cournot equilibrium outcomes. Thus the test of Proposition 2 is only sufficient to reject Cournot behavior in favor of more collusion, e.g. outcome "+" in the figure is consistent not only with constrained Cournot for either firm but also with more (or less) collusive behavior.

4.4.2 The evidence

Prior to testing the supply decisions of firms in geographic space and over time, conditional on plant location, for the entire sample, I consider a specific example extracted from the data. This is reflective of a broader trend, where in many instances Brazilian cement firms undersupply local markets as benchmarked against the supply behavior of a Cournot firm. For further illustrations and details, I refer the reader to Salvo (2006).

A case in point: The supply of 2 firms to 2 markets in 1 time period   Consider the two adjacent states of Alagoas (AL) and Sergipe (SE), located in the northeast of Brazil (see Figure 4). These states are equally small both in terms of market size and geography. Up until 1996 each state was home to only one plant: the firm Brennand

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[51]Following Bernheim and Whinston (1990), the *Supplementary Section* examines the equilibrium support of this type of spatial supply arrangement, where heterogeneous firms meet in multiple markets.

operated the plant located in AL (respectively firm 1 and market A) and its rival Votorantim operated the plant located in SE (respectively firm 2 and market B). Both firms 1 and 2 also owned other plants located in nearby states. Consider the year 1996. While firm 1 commands an 83% share in market A, it chooses not to supply to neighboring market B, right next door to its plant located in market A, despite the large price-cost margin it would enjoy were it to do so. Equally striking, firm 2 commands an 89% share in market B, while attaining only a 7% share in the neighboring market A, next door to its plant in market B. Average consumer prices in markets A and B are almost identical, respectively R\$ 9.46 (per bag) and R\$ 9.44. I calculate firm 1's marginal cost (including sales taxes and the reseller's mark-up) in supplying markets A and B to be respectively R\$ 5.20 and R\$ 5.47. As for firm 2, I calculate its cost in supplying markets A and B to be respectively R\$ 5.30 and R\$ 5.16.[52] This is illustrated in the following picture and table, where I take the price elasticities of demand in equilibrium to be their point estimates (in 1996) from Section 4.2: -0.84 for market AL and -0.18 for market SE.

Market A   Market B

Firm 1   Firm 2
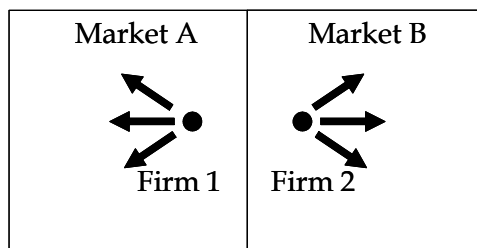
E.g. Market A: AL Market B: SE
      Firm 1: Brennand Firm 2: Votorantim

(Variable means for year 1996)	Price p_{lt}	Share $\frac{q_{flt}}{q_{lt}}$	Marginal Revenue Cournot Point estimate $p_{lt}+\frac{p_{lt}}{\hat{\eta}_{lt}}\frac{q_{flt}}{q_{lt}}$	MC c_{flt}	Reject Cournot?[53] TS $=\hat{\varphi}_{flt,2.5\%}>0$
Local market $l = A$ (AL)					
Firm $f = 1$ (Brennand)	9.46	0.83	$9.46-\frac{9.46}{0.84}0.83 = 0.13$	5.20	No
Firm $f = 2$ (Votoran)	9.46	0.07	$9.46-\frac{9.46}{0.84}0.07 = 8.72$	5.30	**Yes**
Local market $l = B$ (SE)					
Firm $f = 1$ (Brennand)	9.44	0	$9.44-\frac{9.44}{0.18}0 = 9.44$	5.47	**Yes**
Firm $f = 2$ (Votoran)	9.44	0.89	$9.44-\frac{9.44}{0.18}0.89 = -36.38$	5.16	No

I thus reject (given current prices) the hypothesis of firm 2 behaving in Cournot fashion towards market A in 1996, since (perceived) marginal revenue (point estimate of

[52] Note that the state-capital cities of AL (market A) and SE (market B) are located less than 300 km apart. Nevertheless, the difference in Votorantim's (say) cost of supplying AL and SE seems low: only R\$ 0.14. The reason is that Brazil has an awkward sales tax system which may work against within-state shipments, as happens here, i.e. shipments from Votorantim's plant in SE to resellers in SE are penalized compared to its shipments across the state border to resellers in AL. This mitigates the difference in average freight costs from Votorantim's plant in SE: R\$ 0.32 to resellers in SE and R\$ 0.77 to resellers in AL.

[53] When testing the null hypothesis of Cournot behavior, I take the upper bound to the 95% confidence interval for the demand elasticity (which corresponds to the lower bound to the 95% confidence interval for the test statistic φ, since the randomness in $\hat{\varphi}$ stems from the randomness of the estimated demand elasticity $\hat{\eta}$). I return to this below.

8.72) significantly exceeds marginal cost (5.30), with the 95% confidence interval (C.I.) for test statistic $\varphi_{2A,1996}$ being $(2.56, 3.69)$ (and where a point estimate of $8.72 - 5.30 = 3.42$ amounts to $3.42/9.46 = 36\%$ of consumer price). Likewise, I reject Cournot behavior for firm 1 towards market B in 1996 (a point estimate of $\hat{\varphi}_{1B,1996} = 9.44 - 5.47 = 3.97$ amounts to 42% of consumer price!). Notice that firm 2's (resp. firm 1's) supply decision toward market A (resp. market B) corresponds to that of firm f in the outcome marked "•" in Figure 15. A story where Votorantim tacitly agrees to give Brennand the upper hand in AL in exchange for the latter staying away from SE is consistent with observed shipments. Interestingly, Brennand ships from its plant in AL to the states of PB, PE and BA, located at further distances than SE and where prices are similar to those in SE.

Testing the full sample I now present the results for the full sample. From the Cournot pricing condition (18), I compute the test statistic

$$\hat{\varphi}_{flt} = p_{lt} + \frac{p_{lt}}{\hat{\eta}_{lt}} \frac{q_{flt}}{q_{lt}} - c_{flt} \tag{19}$$

for each active-firm-market-month combination (f, l, t), where $\hat{\eta}_{lt}$ is based on the demand estimates of Section 4.2, and p_{lt}, q_{flt} and c_{flt} are observed. A firm is active in a given month if it owns at least one active plant in that month, i.e. firm f is active iff $\sum_l \sum_{i \in \mathcal{O}_{ft}} q_{ilt} > 0$. For every month t in which a firm f is active in the 156 months of the sample, there are 17 (f, l, t) combinations, one for each of the 17 markets, irrespective of the markets to which firm f *actually* ships in month t. There are 37536 active-firm-market-month combinations corresponding, therefore, to an average of $37536/17/156 \approx 14$ active firms in the sample in any given month. For every month t, I take firm f's marginal cost in serving market l, c_{flt}, as the minimum among the marginal costs in serving market l from the plants that it owns, i.e. $c_{flt} := \min_{i \in \mathcal{O}_{ft}} c_{ilt}$; this cost typically corresponds to that of the closest plant (though costs do vary across plants conditional on location on account of other characteristics such as kiln size, kiln age and factor prices). Given the idle capacity that is pervasive in the industry, this reflects firm f's cost of increasing supply to market l on the margin. Recall that, with a view to testing conduct, the marginal costs I construct may conservatively err on the high side (though any systematic bias is small); this would understate $\hat{\varphi}_{flt}$, thus working against the rejection of the null $H_0 : \varphi_{flt} \leq 0$ (constrained Cournot). Finally, I reject the null of Cournot behavior in favor of the alternative of more collusive behavior when the 95% C.I. for the test statistic $\hat{\varphi}_{flt}$ falls entirely on the positive domain, i.e. when the lower bound to the C.I. is greater than zero. Now, since the randomness in $\hat{\varphi}_{flt}$ stems from the randomness of the estimated market demand elasticity $\hat{\eta}_{lt}$, I map the

95% C.I. (lower bound) for $\hat{\varphi}_{flt}$ from the 95% C.I. (upper bound[54]) of $\hat{\eta}_{lt}$. But there is an empirical issue I must overcome: around one-third of the estimated elasticities have an upper bound to the 95% C.I. that crosses over to the positive domain, suggesting that for $\hat{\eta}_{lt}$ at the upper extreme of the interval the demand curve slopes upward! I deal with this here by simply dropping supply decisions that pertain to these market-month pairs (see Salvo 2006 for robustness checks).

Total number of active-firm-market-month combinations, (f, l, t)	37536	
Combinations for which the upper limit to the 95% C.I. for η_{lt} is negative...	24696	100%
and $\hat{\varphi}_{flt}$ is greater than zero...	16806	
and $\hat{\varphi}_{flt}$ is significantly greater than zero at the 5% level	**14849**	60%
and $\hat{\varphi}_{flt}$ is positive and exceeds 10% of consumer price	13197	
and $\hat{\varphi}_{flt}$ is positive and exceeds 20% of consumer price	**8035**	33%

The table above summarizes the results. Of the 24696 active-firm-market-month supply decisions for which the C.I. for the market demand elasticity falls within the interval $(-\infty, 0)$, I find that the null hypothesis of Cournot behavior that allows for the constraining effect of imports, $\varphi_{flt} \leq 0$, can be *rejected at the 5% level of significance in 14849 instances*. In other words, under the Cournot conjecture, one would expect firms to expand their supply to local markets in $14849/24696 \simeq 60\%$ of supply decisions vis-à-vis observed shares – these firms are choosing output to the left of their Cournot reaction functions. As in the earlier illustration, the test statistics $\hat{\varphi}_{flt}$ are not only positive but sizeable: the point estimate for $\hat{\varphi}_{flt}$ exceeds 20% of consumer price in 8035 supply decisions! Firms are typically serving their own home turf and restricting supply to other markets, no matter how profitable this would be in a static sense.

5 Concluding remarks

This paper shows that the standard methodology for inferring supply can underestimate the degree of market power in a setting where firms are constrained, in ways unobserved to the analyst, in their ability to set prices in response to changing demand conditions. The finding is of high practical relevance on two counts. First, the methodology has been used widely by both academics and practitioners seeking to quantify market power in the absence of cost data. Second, one can argue that the extent to which unobserved constraints impinge on market outcomes is empirically important.

In the paper, I model a domestic oligopoly that is constrained by the potential entry of high-cost imports. In so doing, I can naturally refer to the case of the Brazilian cement

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[54]To see this, notice from (19) that increasing (a negative) $\hat{\eta}$ toward zero lowers $\hat{\varphi}$.

industry, which provides a clear-cut illustration of my methodological result. This choice of model and its empirical illustration also suggest that the result is increasingly relevant in a world where trade barriers are being pulled down and geographic markets are coming closer together. More generally, one can conceive of several other dynamic channels by which constraints on the ability to price may operate. Examples range from the threat of antitrust litigation or regulatory intervention, to entry games with hysteresis or signaling, where unobserved price thresholds may be motivated.

This result recommends caution to enthusiasts of the measurement of market power when cost data is lacking. Unless the analyst is somehow sure that the data is *not* constrained in ways not already captured by the structural model (typically the behavior of consumers and existing competitors), his estimated price-cost margins may potentially be only lower bounds to the true margins. The paper suggests that a tall order – the identification of supply – is, in practice, even taller.

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A   Appendix: (Sketch of) Proof of Corollary 1

I begin by motivating a simple structural model by reference to a domestic cement industry facing latent high-cost import competition. I then analytically derive $\left(\text{plim } \hat{\theta} - \theta \right) < 0$ for the simple case. Finally, to demonstrate the generality of this analytical result, I report results for a fraction of the Montecarlo experiments that I have conducted under a wide range of demand and supply specifications. Results for further specifications are available upon request (e.g. different functional forms for demand and for domestic marginal cost, varying the covariance matrix of the exogenous variables).

A.1   A simple motivating example: linear demand and linear cost

In a given local market (for cement, say), demand (1) 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) $\varepsilon^d \sim N(0, \sigma_d^2)$. The 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) $W_1 \sim N(\bar{W}_1, \sigma_{W_1}^2)$ 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) $W_2 \sim N(\bar{W}_2, \sigma_{W_2}^2)$ 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 of the exogenous variables, where (i) $\beta_4 > 0$ reflects the (energy) cost of international transportation; and (ii) $W_3 = \frac{W_2}{e} + T$ is the "world" price of labor, in US\$, and $T \sim N(\bar{T}, \sigma_T^2)$ 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. Complete the structural model with the DGP (4), where $\varepsilon^s \sim N(0, \sigma_s^2)$, noting that, in view of the domestic monopoly, $\theta = 1$.

A.2 Analytical proof of Corollary 1 for the simple motivating example, with just identification $J = K = 2$

Simplify the motivating example further by considering the special case where domestic marginal cost is

$$c(W; \beta) = \beta W$$

and the marginal cost of imports is

$$c^I(W, T; \beta) = \beta(W + T)$$

(All exogenous covariates Y_1, Y_2 – on the demand side – and W, T – on the supply side – are assumed to be orthogonal to one another.) The estimated model (5) is

$$p = X\delta + \xi^s \tag{20}$$

where, recalling the notation of the proof of Proposition 1, $X = (X_1 \ \ W)$, $X_1 = -q\frac{\partial p(q,.)}{\partial q} = -q(\alpha_2 + \alpha_4 Y_2)^{-1} > 0$ (linear demand), and $\delta = (\theta \ \ \beta)'$. Consider the matrix of instruments $Z = (Y_1 \ \ W)$ where $rank(Z) = rank(X) = 2$. The 2SLS estimator (11)

collapses to

$$\hat{\delta} = (Z'X)^{-1}Z'p$$

which rearranges to

$$
\begin{aligned}
Z'X\hat{\delta} &= Z'p \\
(Y_1 \ \ W)'(X_1 \ \ W) \begin{pmatrix} \hat{\theta} \\ \hat{\beta} \end{pmatrix} &= (Y_1 \ \ W)'p
\end{aligned}
$$

or the system of equations

$$
\left.
\begin{aligned}
Y_1'X_1\hat{\theta} + Y_1'W\hat{\beta} &= Y_1'p \\
W'X_1\hat{\theta} + W'W\hat{\beta} &= W'p
\end{aligned}
\right\}
$$

Solving for $\hat{\theta}$ (noting that $W'W$ and $(W'WY_1' - Y_1'WW')X_1$ are scalars), and substituting for p from (20):

$$
\begin{aligned}
\hat{\theta} &= \frac{(W'WY_1' - Y_1'WW')\,p}{(W'WY_1' - Y_1'WW')\,X_1} \\
&= \theta + \frac{(W'WY_1' - Y_1'WW')\,W\beta}{(W'WY_1' - Y_1'WW')\,X_1} + \frac{(W'WY_1' - Y_1'WW')\,\xi^s}{(W'WY_1' - Y_1'WW')\,X_1} \\
&= \theta + \frac{\left(\frac{1}{N}W'W\right)\left(\frac{1}{N}Y_1'\xi^s\right) - \left(\frac{1}{N}Y_1'W\right)\left(\frac{1}{N}W'\xi^s\right)}{\left(\frac{1}{N}W'W\right)\left(\frac{1}{N}Y_1'X_1\right) - \left(\frac{1}{N}Y_1'W\right)\left(\frac{1}{N}W'X_1\right)}
\end{aligned}
$$

since $(W'WY_1' - Y_1'WW')\,W = 0$. Applying a standard law of large numbers to each term along with Slutsky's theorem yields

$$
\begin{aligned}
\operatorname{plim}\hat{\theta} &= \theta + \operatorname{plim}\left(\frac{\left(\frac{1}{N}W'W\right)\left(\frac{1}{N}Y_1'\xi^s\right) - \left(\frac{1}{N}Y_1'W\right)\left(\frac{1}{N}W'\xi^s\right)}{\left(\frac{1}{N}W'W\right)\left(\frac{1}{N}Y_1'X_1\right) - \left(\frac{1}{N}Y_1'W\right)\left(\frac{1}{N}W'X_1\right)} \right) \qquad (21) \\
&= \theta + \frac{E\left(W^2\right)E\left(Y_1\xi^s\right) - E\left(Y_1W\right)E\left(W\xi^s\right)}{E\left(W^2\right)E\left(Y_1X_1\right) - E\left(Y_1W\right)E\left(WX_1\right)} \\
&= \theta + \frac{Var\left(W\right)E\left(Y_1\right)E\left(\xi^s\right) + E\left(W^2\right)Cov\left(Y_1,\xi^s\right) - E\left(Y_1\right)E\left(W\right)Cov\left(W,\xi^s\right)}{Var\left(W\right)E\left(Y_1\right)E\left(X_1\right) + E\left(W^2\right)Cov\left(Y_1,X_1\right) - E\left(Y_1\right)E\left(W\right)Cov\left(W,X_1\right)}
\end{aligned}
$$

noting that $Cov\left(Y_1,W\right) = 0$ (and where letters now denote random variables rather than vectors of observed or calculated covariates).

The remainder of the proof consists in showing that, for all realizations of the exogenous variables, the denominator of the second term of (21) is positive, while the numerator is negative, such that $\operatorname{plim}\hat{\theta} < \theta$ obtains. Begin by considering

$$
\begin{aligned}
X_1 &= -\left(\alpha_2 + \alpha_4 Y_2\right)^{-1}\left(q^*\Pr(\chi = 1) + q\Pr(\chi = 0)\right) \\
&= \frac{\alpha_1 + \alpha_3 Y_1 + \varepsilon^d + \left(\alpha_2 + \alpha_4 Y_2\right)(\beta W + \varepsilon^s)}{-\left(\alpha_2 + \alpha_4 Y_2\right)\left(1 + \theta\right)}\Pr(\chi = 1) + \frac{\alpha_1 + \alpha_3 Y_1 + \varepsilon^d + \left(\alpha_2 + \alpha_4 Y_2\right)\beta(W + T)}{-\left(\alpha_2 + \alpha_4 Y_2\right)}\Pr(\chi = 0)
\end{aligned}
$$

where $\Pr(\chi = 1) = \Pr(p^* \leq c^I) = \Pr\left(\varepsilon^s \leq \theta\frac{\alpha_1 + \alpha_3 Y_1 + \varepsilon^d + (\alpha_2 + \alpha_4 Y_2)(\beta W + \varepsilon^s)}{(\alpha_2 + \alpha_4 Y_2)(1+\theta)} + \beta T\right)$. We wish to determine $Cov(Y_1, X_1)$. Conditional on $\chi = 1$,

$$E(Y_1 X_1) = \frac{1}{(1+\theta)} E\left(\frac{Y_1(\alpha_1 + \varepsilon^d)}{-(\alpha_2 + \alpha_4 Y_2)}\right) + \frac{\alpha_3}{(1+\theta)} E\left(\frac{Y_1^2}{-(\alpha_2 + \alpha_4 Y_2)}\right) - \frac{1}{(1+\theta)} E(Y_1(\beta W + \varepsilon^s))$$

and, given that the (conditional) covariance matrix of the exogenous variables is (still) diagonal, we have

$$\begin{aligned} Cov(Y_1, X_1) &= E(Y_1 X_1) - E(Y_1) E(X_1) \\ &= \frac{\alpha_3}{(1+\theta)} E\left(\frac{1}{-(\alpha_2 + \alpha_4 Y_2)}\right) Var(Y_1) > 0 \end{aligned}$$

Similarly, conditional on $\chi = 0$,

$$E(Y_1 X_1) = E\left(\frac{Y_1(\alpha_1 + \varepsilon^d)}{-(\alpha_2 + \alpha_4 Y_2)}\right) + \alpha_3 E\left(\frac{Y_1^2}{-(\alpha_2 + \alpha_4 Y_2)}\right) - \beta E(Y_1(W + T))$$

and

$$Cov(Y_1, X_1) = \alpha_3 E\left(\frac{1}{-(\alpha_2 + \alpha_4 Y_2)}\right) Var(Y_1) > 0$$

Finally, since covariance is a linear operator

$$Cov(Y_1, X_1) = Cov(Y_1, X_1 \mid \chi = 1) \Pr(\chi = 1) + Cov(Y_1, X_1 \mid \chi = 0) \Pr(\chi = 0) > 0$$

We now wish to determine $Cov(W, X_1)$. Conditional on $\chi = 1$,

$$E(W X_1) = \frac{1}{(1+\theta)} E\left(\frac{W(\alpha_1 + \alpha_3 Y_1 + \varepsilon^d)}{-(\alpha_2 + \alpha_4 Y_2)}\right) - \frac{\beta}{(1+\theta)} E(W^2) - \frac{1}{(1+\theta)} E(W\varepsilon^s)$$

and

$$\begin{aligned} Cov(W, X_1) &= E(W X_1) - E(W) E(X_1) \\ &= \frac{-\beta}{(1+\theta)} Var(W) < 0 \end{aligned}$$

Conditional on $\chi = 0$,

$$E(W X_1) = E\left(\frac{W(\alpha_1 + \alpha_3 Y_1 + \varepsilon^d)}{-(\alpha_2 + \alpha_4 Y_2)}\right) - \beta E(W^2) - \beta E(WT)$$

and

$$Cov(W, X_1) = -\beta Var(W) < 0$$

which makes

$$Cov\left(W, X_1\right) = Cov\left(W, X_1 \mid \chi = 1\right)\Pr(\chi = 1) + Cov\left(W, X_1 \mid \chi = 0\right)\Pr(\chi = 0) < 0$$

and therefore the denominator of the second term of (21) is positive (since its first two terms are positive and its third term is negative).

Now consider the error of the estimated model

$$
\begin{aligned}
\xi^s &= \varepsilon^s \Pr(\chi = 1) + (c_I - \theta X_1 - c)\Pr(\chi = 0) \\
&= \varepsilon^s \Pr(\chi = 1) + (\beta\theta W + \beta(1 + \theta)T - \theta\left(\frac{\alpha_1 + \alpha_3 Y_1 + \varepsilon^d}{-(\alpha_2 + \alpha_4 Y_2)}\right))\Pr(\chi = 0)
\end{aligned}
$$

We wish to determine $Cov\left(Y_1, \xi^s\right)$. Conditional on $\chi = 1$,

$$Cov\left(Y_1, \xi^s\right) = Cov\left(Y_1, \varepsilon^s\right) = 0$$

Conditional on $\chi = 0$,

$$E\left(Y_1 \xi^s\right) = E\left(Y_1\left(\beta\theta W + \beta(1 + \theta)T - \theta\left(\frac{\alpha_1 + \varepsilon^d}{-(\alpha_2 + \alpha_4 Y_2)}\right)\right)\right) - \alpha_3\theta E\left(\frac{Y_1^2}{-(\alpha_2 + \alpha_4 Y_2)}\right)$$

and, again, since the exogenous variables are conditionally orthogonal

$$Cov\left(Y_1, \xi^s\right) = -\alpha_3\theta E\left(\frac{1}{-(\alpha_2 + \alpha_4 Y_2)}\right)Var\left(Y_1\right) < 0$$

The linearity of the covariance operator implies that the unconditional covariance

$$Cov\left(Y_1, \xi^s\right) = Cov\left(Y_1, \xi^s \mid \chi = 1\right)\Pr(\chi = 1) + Cov\left(Y_1, \xi^s \mid \chi = 0\right)\Pr(\chi = 0) < 0$$

We now wish to determine $Cov\left(W, \xi^s\right)$. Conditional on $\chi = 1$,

$$Cov\left(W, \xi^s\right) = Cov\left(W, \varepsilon^s\right) = 0$$

Conditional on $\chi = 0$,

$$E\left(W\xi^s\right) = E\left(W\left(\beta(1 + \theta)T - \theta\left(\frac{\alpha_1 + \alpha_3 Y_1 + \varepsilon^d}{-(\alpha_2 + \alpha_4 Y_2)}\right)\right)\right) + \beta\theta E\left(W^2\right)$$

and thus

$$Cov\left(W, \xi^s\right) = \beta\theta Var\left(W\right) > 0$$

so that the unconditional covariance $Cov\left(W, \xi^s\right) > 0$. Therefore the numerator of the second term of (21) is negative (since its first two terms are negative – recall from (8) that $E\left(\xi^s\right) < 0$ – and its third term is positive). $\blacksquare$

A.3 Montecarlo results for some generalizations

Again, the DGP is given by (4), the estimated model is given by (5), and in each experiment $c^I > c(D(c^I, .), .)$ holds for all realizations of the exogenous variables. Denote the number of simulations by S (index each simulation by s) and denote the number of observations generated for each simulation s by N (index each observation by i).

A.3.1 Experiment 1: Linear demand and flat marginal cost

This simulates the simple motivating example of Section A.1. I pick $S = 2000$ and $N = 1000$. Demand is specified as

$$
\begin{aligned}
q &= \alpha_1 + \alpha_2 p + \alpha_3 Y_1 + \alpha_4 p Y_2 + \varepsilon^d \\
Y_1 &\sim N(20, 1^2), Y_2 \sim N(1, .1^2), \varepsilon^d \sim N(0, 1^2) \\
\alpha_1 &= 10, \alpha_2 = -1, \alpha_3 = 1, \alpha_4 = -.2
\end{aligned}
$$

such that $-q \partial p(q, .)/\partial q$ becomes $-q (\alpha_2 + \alpha_4 Y_2)^{-1}$. Supply is specified as

$$
\begin{aligned}
c &= \beta_1 W_1 e + \beta_2 W_2 \\
c^I &= \beta_1 (1 + \beta_4) W_1 e + \beta_3 W_3 e \\
W_1 &\sim N(3, .3^2), W_2 \sim N(3, .3^2), e \sim N(1, .1^2), W_3 = \frac{W_2}{e} + T, T \sim N(3, .3^2), \varepsilon^s \sim N(0, 1^2) \\
\beta_1 &= 1, \beta_2 = 1, \beta_3 = 1, \beta_4 = 2.5, \theta = 1
\end{aligned}
$$

The reduced-form solution to the constraint-free system is given by

$$
(1 + \theta) q^* = \alpha_1 + \alpha_3 Y_1 + \varepsilon^d + (\alpha_2 + \alpha_4 Y_2)(\beta_1 W_1 e + \beta_2 W_2 + \varepsilon^s)
$$

Results Over simulations s, the proportion of constrained observations i (where $p_i = c_i^I$) ranges from 0.227 **to** 0.338. Demand estimates using the observed variables p, q appear consistent (if anything, the greater variation vis-à-vis estimation with the latent variables p^*, q^* seems to lead to greater efficiency). On the supply side, while estimation with latent variables is consistent (Bresnahan 1982), *estimation with observed variables p, q understates the true degree of market power in all $S = 2000$ simulations.* I incorrectly reject the hypothesis that $\theta = 1$ in favor of greater competition in all $S = 2000$ simulations. The estimated elasticity-adjusted Lerner index averages (i) 57% across unconstrained observations (against a true index of 99%), and (ii) 56% across constrained observations (against a true index of 86%). Cost is overestimated.

DEMAND		
Estimation with:	**Observed** variables p, q	**Latent** variables p^*, q^*
	Full sample $\chi_i \in \{0, 1\}$	Full sample $\chi_i \in \{0, 1\}$
Mean (Std.Dev.) over simulations s		
$\hat{\alpha}_{1,s}$ (True $\alpha_1 = 10$)	9.963 (0.834)	9.826 (1.113)
$\hat{\alpha}_{2,s}$ (True $\alpha_2 = -1$)	-0.997 (0.057)	-0.980 (0.113)
$\hat{\alpha}_{3,s}$ (True $\alpha_3 = 1$)	0.999 (0.036)	0.990 (0.062)
$\hat{\alpha}_{4,s}$ (True $\alpha_4 = -.2$)	-0.199 (0.022)	-0.197 (0.027)

Notes to the table: (1) Excluded exogenous variables serving as instruments for endogenous variables p and pY_2 are: W_1, W_2, W_3, and interactions of these covariates with Y_2. (2) Specifying just identification (e.g. using only W_1 and W_1Y_2 as instruments) yields similar estimates. (3) Interacting instruments with e (say W_1e) also yields similar estimates.

SUPPLY	**Threat of entry**	**Bresnahan (1982)**
Estimation with:	**Observed** variables p, q	**Latent** variables p^*, q^*
	Full sample $\chi_i \in \{0, 1\}$	Full sample $\chi_i \in \{0, 1\}$
Mean (Std.Dev.) over simulations s		
$\hat{\theta}_s$ (True $\theta = 1$)	**0.564 (0.025)**	**0.999 (0.029)**
$\hat{\beta}_{1,s}$ (True $\beta_1 = 1$)	2.204 (0.071)	1.002 (0.067)
$\hat{\beta}_{2,s}$ (True $\beta_2 = 1$)	1.023 (0.082)	1.001 (0.090)

Count $s \in \{1, ..., 2000\}$ such that the 95% C.I. for $\hat{\theta}_s^{\mathbf{observed}}$ lies below $\theta^{\mathbf{true}}$: **2000** (i.e. 100%)

Count $s \in \{1, ..., 2000\}$ such that the 95% C.I. for $\hat{\theta}_s^{\mathbf{latent}}$ lies below $\theta^{\mathbf{true}}$: **63** (i.e. 3.2%)

Further detail of estimation with **observed** variables p, q and full sample:

Mean (Std.Dev.) over simulations s where, for every s, take the mean over observations i...

	True, c:	Estimate, $\hat{c}$:	True, $\eta\frac{p-c}{p}$:	Estimate, $\eta\frac{p-\hat{c}}{p}$:
...such that $\chi_i = 1$	6.182 (0.017)	10.066 (0.248)	**0.991 (0.004)**	**0.573 (0.026)**
	True, c:	Estimate, $\hat{c}$:	True, $\eta\frac{p-c}{p}$:	Estimate, $\eta\frac{p-\hat{c}}{p}$:
...such that $\chi_i = 0$	5.548 (0.023)	8.720 (0.225)	**0.856 (0.008)**	**0.559 (0.024)**

Notes to the table: (1) Excluded exogenous variables serving as instruments for endogenous variable $-q\partial p(q, .)/\partial q = -q\left(\alpha_2 + \alpha_4Y_2\right)^{-1}$ are: Y_1, Y_2 (very similar estimates are obtained if $Y_1\left(\alpha_2 + \alpha_4Y_2\right)^{-1}$, $Y_2\left(\alpha_2 + \alpha_4Y_2\right)^{-1}$ are used instead). (2) Similar estimates are obtained if a constant is added to the estimated model. (3) For every simulation s, evaluate true c and estimated $\hat{c}$ at the mean value of covariates across i. (4) For every simulation s, evaluate the true and estimated elasticity-adjusted Lerner index for each observation i and then take mean values across i (estimate assumes η is known to abstract away from sampling error and focus on the supply inconsistency).

A.3.2 Experiment 2: Linear demand and linear marginal cost, non-diagonal covariance matrix

Domestic marginal cost is now linearly increasing in output. The covariance matrix of the exogenous variables is no longer diagonal: $Corr(Y_1, Y_2) > 0$ and $Corr(Y_1, W_2) > 0$ as specified below. I maintain $S = 2000$ and $N = 1000$. Demand is specified as

$$
\begin{aligned}
q &= \alpha_1 + \alpha_2 p + \alpha_3 Y_1 + \alpha_4 p Y_2 + \varepsilon^d \\
Y_1 &\sim N(20, 1^2), Y_2 = \frac{Y_1}{20}\hat{Y}_2, \hat{Y}_2 \sim N(1, .1^2), \varepsilon^d \sim N(0, 1^2) \\
\alpha_1 &= 10, \alpha_2 = -1, \alpha_3 = 1, \alpha_4 = -.2
\end{aligned}
$$

such that $-q\partial p(q, .)/\partial q$ is $-q(\alpha_2 + \alpha_4 Y_2)^{-1}$. Supply is now specified as

$$
\begin{aligned}
c &= \beta_1 W_1 e + \beta_2 W_2 q \\
c^I &= \beta_1(1 + \beta_4)W_1 e + \beta_3 W_3 e \\
W_1 &\sim N(3, .3^2), W_2 = \frac{Y_1}{20}\hat{W}_2, \hat{W}_2 \sim N(.3, .03^2), e \sim N(1, .1^2), W_3 = 10\frac{W_2}{e} + T, \\
T &\sim N(3, .3^2), \varepsilon^s \sim N(0, 1^2) \\
\beta_1 &= 1, \beta_2 = 1, \beta_3 = 1, \beta_4 = 2.5, \theta = 1
\end{aligned}
$$

The reduced-form solution to the constraint-free system is given by

$$
(1 + \theta - (\alpha_2 + \alpha_4 Y_2)\beta_2 W_2) q^* = \alpha_1 + \alpha_3 Y_1 + \varepsilon^d + (\alpha_2 + \alpha_4 Y_2)(\beta_1 W_1 e + \varepsilon^s)
$$

Results Over simulations s, the proportion of constrained observations i (where $p_i = c_i^I$) ranges from 0.260 **to** 0.368. Similar comments to those in Experiment 1 above apply. Importantly, *estimation with observed variables p, q again understates the true degree of market power in all $S = 2000$ simulations*. (I again incorrectly reject the hypothesis that $\theta = 1$ in favor of greater competition in all $S = 2000$ simulations.)

DEMAND		
Estimation with:	**Observed** variables p, q	**Latent** variables p^*, q^*
	Full sample $\chi_i \in \{0, 1\}$	Full sample $\chi_i \in \{0, 1\}$
Mean (Std.Dev.) over simulations s		
$\hat{\alpha}_{1,s}$ (True $\alpha_1 = 10$)	9.976 (0.804)	9.842 (0.960)
$\hat{\alpha}_{2,s}$ (True $\alpha_2 = -1$)	-0.997 (0.055)	-0.971 (0.125)
$\hat{\alpha}_{3,s}$ (True $\alpha_3 = 1$)	0.998 (0.042)	0.981 (0.086)
$\hat{\alpha}_{4,s}$ (True $\alpha_4 = -.2$)	-0.199 (0.021)	-0.195 (0.028)

Notes to the table: (1) Excluded exogenous variables serving as instruments for endogenous variables p and pY_2 are: W_1, W_2, W_3, and interactions of these covariates with Y_2. (2)

Specifying just identification (e.g. using only W_1 and W_1Y_2 as instruments) yields similar estimates. (3) Interacting instruments with e (say W_1e) also yields similar estimates.

SUPPLY	**Threat of entry**	**Bresnahan (1982)**
Estimation with:	**Observed** variables p, q	**Latent** variables p^*, q^*
	Full sample $\chi_i \in \{0, 1\}$	Full sample $\chi_i \in \{0, 1\}$
Mean (Std.Dev.) over simulations s		
$\hat{\theta}_s$ (True $\theta = 1$)	**0.437** **(0.036)**	**0.999** **(0.036)**
$\hat{\beta}_{1,s}$ (True $\beta_1 = 1$)	2.642 (0.073)	1.003 (0.066)
$\hat{\beta}_{2,s}$ (True $\beta_2 = 1$)	0.913 (0.079)	1.000 (0.082)
Count $s \in \{1, ..., 2000\}$ such that the 95% C.I. for $\hat{\theta}_s^{\textbf{observed}}$ lies below $\theta^{\textbf{true}}$: **2000** (i.e. 100%)		
Count $s \in \{1, ..., 2000\}$ such that the 95% C.I. for $\hat{\theta}_s^{\textbf{latent}}$ lies below $\theta^{\textbf{true}}$: **63** (i.e. 3.2%)		

Further detail of estimation with **observed** variables p, q and full sample:

Mean (Std.Dev.) over simulations s where, for every s, take the mean over observations i...

	True, c:	Estimate, $\hat{c}$:	True, $\eta\frac{p-c}{p}$:	Estimate, $\eta\frac{p-\hat{c}}{p}$:
...such that $\chi_i = 1$	6.509 (0.020)	11.461 (0.335)	**0.991** **(0.004)**	**0.449** **(0.036)**
	True, c:	Estimate, $\hat{c}$:	True, $\eta\frac{p-c}{p}$:	Estimate, $\eta\frac{p-\hat{c}}{p}$:
...such that $\chi_i = 0$	6.423 (0.031)	10.348 (0.343)	**0.803** **(0.009)**	**0.428** **(0.035)**

Notes to the table: (1) Excluded exogenous variables serving as instruments for endogenous variables $-q\partial p(q, .)/\partial q = -q\left(\alpha_2 + \alpha_4Y_2\right)^{-1}$ and W_2q are: Y_1, Y_2 (very similar estimates are obtained if $Y_1\left(\alpha_2 + \alpha_4Y_2\right)^{-1}$, $Y_2\left(\alpha_2 + \alpha_4Y_2\right)^{-1}$ are used instead), and W_2Y_1, W_2Y_2. (2) Specifying just identification (e.g. using only Y_2 and W_2Y_1 as instruments) yields similar estimates. (3) Similar estimates are obtained if a constant is added to the estimated model. (4) For every simulation s, evaluate the true and estimated marginal cost (c and $\hat{c}$) for each observation i and then take mean values across i. (5) For every simulation s, evaluate the true and estimated elasticity-adjusted Lerner index for each observation i and then take mean values across i (estimate assumes η is known to abstract away from sampling error and focus on the supply inconsistency).

A.3.3   Experiment 3: Exponential demand and linear marginal cost

I now specify exponential demand. I lower the number of simulations to $S = 500$ (to reduce computational time, given that the system is now non-linear – see below) and keep $N = 1000$. Demand is

$$
\begin{aligned}
\ln(q) &= \alpha_1 + \alpha_2 p + \alpha_3 Y_1 + \alpha_4 pY_2 + \varepsilon^d \\
Y_1 &\sim N(10, 1^2), Y_2 \sim N(3, .6^2), \varepsilon^d \sim N(0, .4^2) \\
\alpha_1 &= -.2, \alpha_2 = -.1, \alpha_3 = .3, \alpha_4 = -.02
\end{aligned}
$$

such that $-q\partial p(q,.)/\partial q$ becomes $-\left(\alpha_2 + \alpha_4 Y_2\right)^{-1}$. Supply is now specified as

$$
\begin{aligned}
c &= \beta_1 W_1 e + \beta_2 W_2 q \\
c^I &= \beta_1 (1 + \beta_4) W_1 e + \beta_3 W_3 e \\
W_1 &\sim N(3, .3^2), W_2 \sim N(.2, .02^2), e \sim N(1, .1^2), W_3 = 5\frac{W_2}{e} + T, T \sim N(4, .4^2), \varepsilon^s \sim N(0, 1^2) \\
\beta_1 &= 1, \beta_2 = 1, \beta_3 = 1, \beta_4 = 1, \theta = 1
\end{aligned}
$$

The reduced-form solution to the constraint-free system is obtained implicitly:

$$
\ln(q^*) - \left(\alpha_2 + \alpha_4 Y_2\right)\beta_2 W_2 q^* = \alpha_1 + \alpha_3 Y_1 + \varepsilon^d - \theta + \left(\alpha_2 + \alpha_4 Y_2\right)\left(\beta_1 W_1 e + \varepsilon^s\right)
$$

Results Over simulations s, the proportion of constrained observations i (where $p_i = c_i^I$) ranges from 0.203 **to** 0.289. Similar comments to those for the experiments above apply. Importantly, *estimation with observed variables p, q again understates the true degree of market power in all $S = 500$ simulations.* (I again incorrectly reject the hypothesis that $\theta = 1$ in favor of greater competition in all $S = 500$ simulations.)

DEMAND		
Estimation with:	**Observed** variables p, q	**Latent** variables p^*, q^*
	Full sample $\chi_i \in \{0, 1\}$	Full sample $\chi_i \in \{0, 1\}$
Mean (Std.Dev.) over simulations s		
$\hat{\alpha}_{1,s}$ (True $\alpha_1 = -.2$)	-0.243 (0.361)	-0.227 (0.311)
$\hat{\alpha}_{2,s}$ (True $\alpha_2 = -.1$)	-0.096 (0.032)	-0.098 (0.028)
$\hat{\alpha}_{3,s}$ (True $\alpha_3 = .3$)	0.299 (0.015)	0.300 (0.014)
$\hat{\alpha}_{4,s}$ (True $\alpha_4 = -.02$)	-0.020 (0.004)	-0.020 (0.003)

Notes to the table: (1) Excluded exogenous variables serving as instruments for endogenous variables p and pY_2 are: W_1, W_2, W_3, and interactions of these covariates with Y_2. (2) Specifying just identification (e.g. using only W_1 and $W_1 Y_2$ as instruments) yields similar estimates. (3) Interacting instruments with e (say $W_1 e$) also yields similar estimates.

SUPPLY	Threat of entry	Bresnahan (1982)
Estimation with:	**Observed** variables p, q	**Latent** variables p^*, q^*
	Full sample $\chi_i \in \{0, 1\}$	Full sample $\chi_i \in \{0, 1\}$

Mean (Std.Dev.) over simulations s

$\hat{\theta}_s$ (True $\theta = 1$)	**0.746 (0.039)**	**0.998 (0.041)**
$\hat{\beta}_{1,s}$ (True $\beta_1 = 1$)	1.489 (0.065)	1.004 (0.069)
$\hat{\beta}_{2,s}$ (True $\beta_2 = 1$)	0.860 (0.147)	1.002 (0.176)

Count $s \in \{1, ..., 500\}$ such that the 95% C.I. for $\hat{\theta}_s^{\mathbf{observed}}$ lies below $\theta^{\mathbf{true}}$: **500** (i.e. 100%)

Count $s \in \{1, ..., 500\}$ such that the 95% C.I. for $\hat{\theta}_s^{\mathbf{latent}}$ lies below $\theta^{\mathbf{true}}$: **19** (i.e. 3.8%)

Further detail of estimation with **observed** variables p, q and full sample:

Mean (Std.Dev.) over simulations s where, for every s, take the mean over observations i...

	True, c:	Estimate, $\hat{c}$:	True, $\eta\frac{p-c}{p}$:	Estimate, $\eta\frac{p-\hat{c}}{p}$:
...such that $\chi_i = 1$	3.811 (0.019)	5.222 (0.244)	**0.958 (0.006)**	**0.730 (0.038)**
...such that $\chi_i = 0$	True, c: 3.650 (0.038)	Estimate, $\hat{c}$: 4.849 (0.250)	True, $\eta\frac{p-c}{p}$: **0.983 (0.009)**	Estimate, $\eta\frac{p-\hat{c}}{p}$: **0.799 (0.040)**

Notes to the table: (1) Excluded exogenous variable serving as instrument for endogenous variable $W_2 q$ is $W_2 Y_1$ (notice that $X_1 = -q \partial p(q, .)/\partial q = -(\alpha_2 + \alpha_4 Y_2)^{-1}$ is exogenous). (2) Similar estimates are obtained if a constant is added to the estimated model. (3) For every simulation s, evaluate the true and estimated marginal cost (c and $\hat{c}$) for each observation i and then take mean values across i. (4) For every simulation s, evaluate the true and estimated elasticity-adjusted Lerner index for each observation i and then take mean values across i (estimate assumes η is known to abstract away from sampling error and focus on the supply inconsistency).
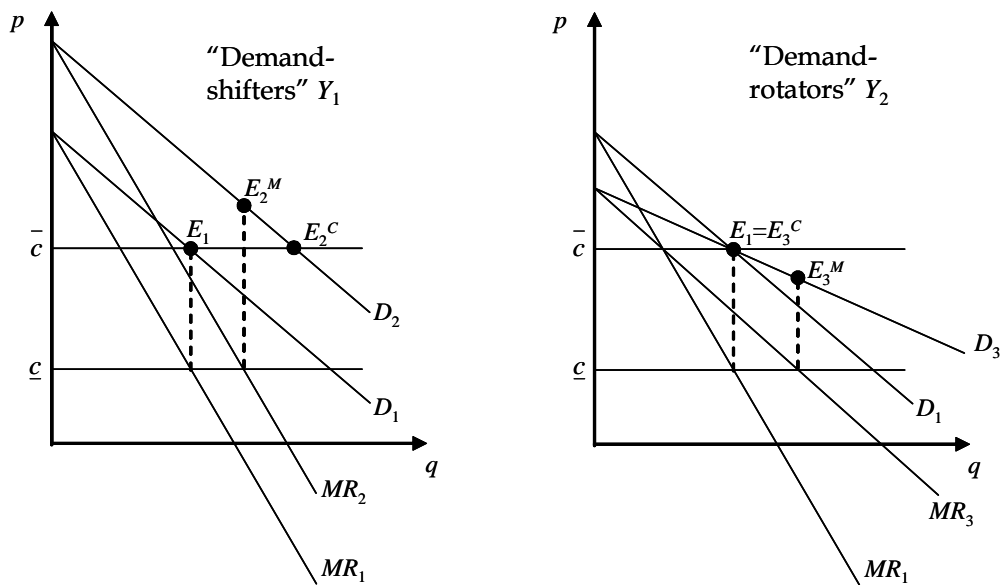
Figure 1: Identification in a static model (for the linear demand and linear cost example).
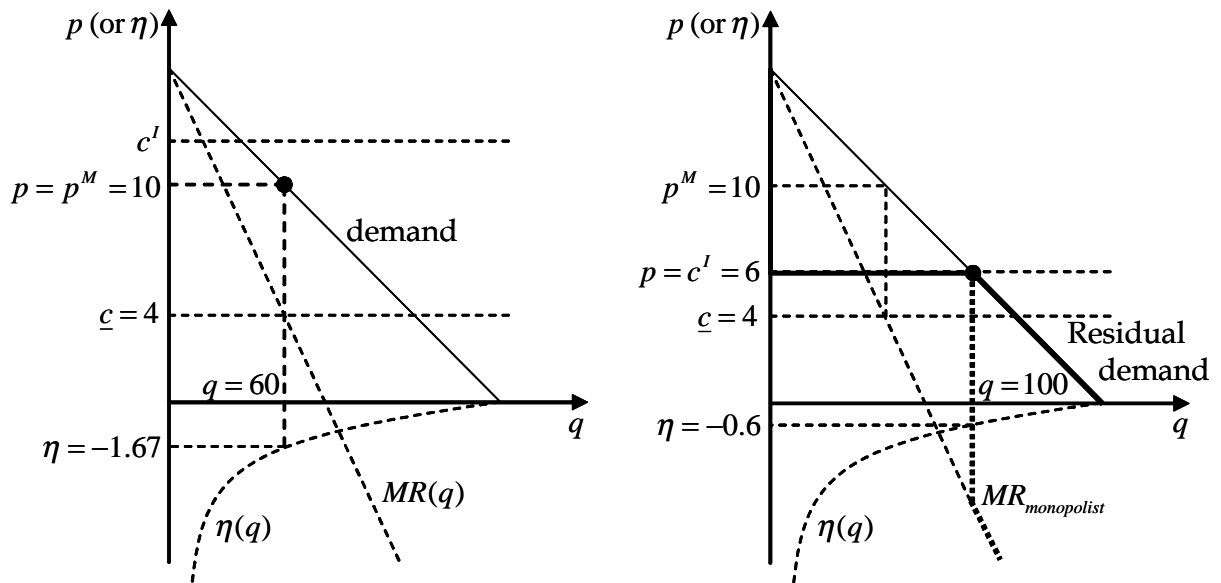Left panel: Demand shifts. Right panel: Demand rotates.

Figure 2: Monopolist facing a competitive fringe. Drawn for linear demand and linear cost, given by $p = 16 - \frac{1}{10}q$, and $\underline{c} = 4$. (The market price elasticity of demand as a function of q is then $\eta(q) = 1 - 160q^{-1}$.) Left panel: Imports have no "bite" ($p^M \leq c^I$). Right panel: Imports constrain price in equilibrium ($p^M > c^I = 6$, as drawn).
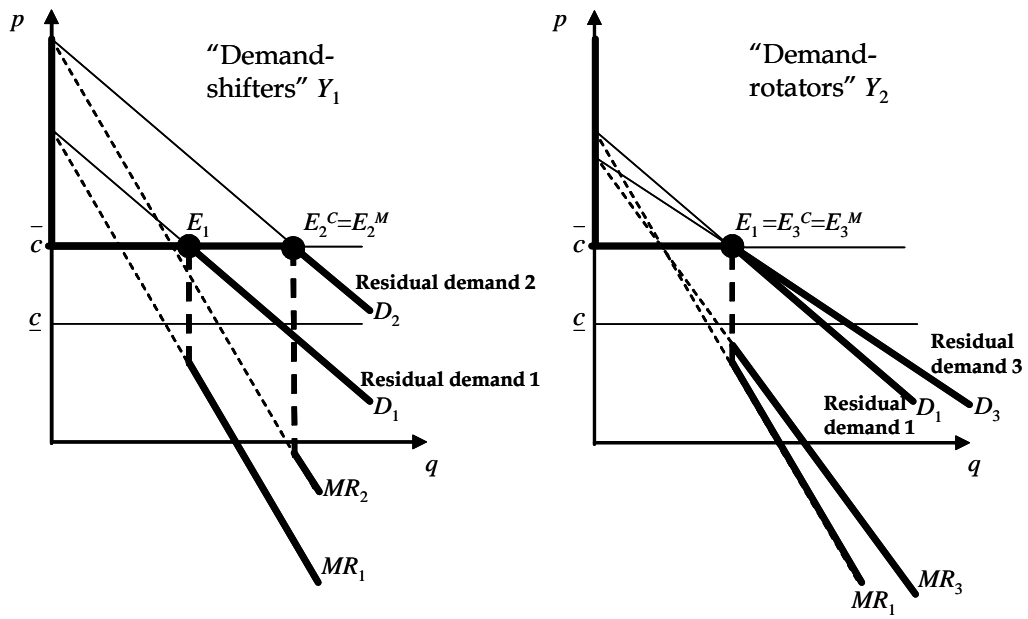
Figure 3: Conduct is no longer identified under the threat of entry. Left panel: Demand shifts. Right panel: Demand rotates.
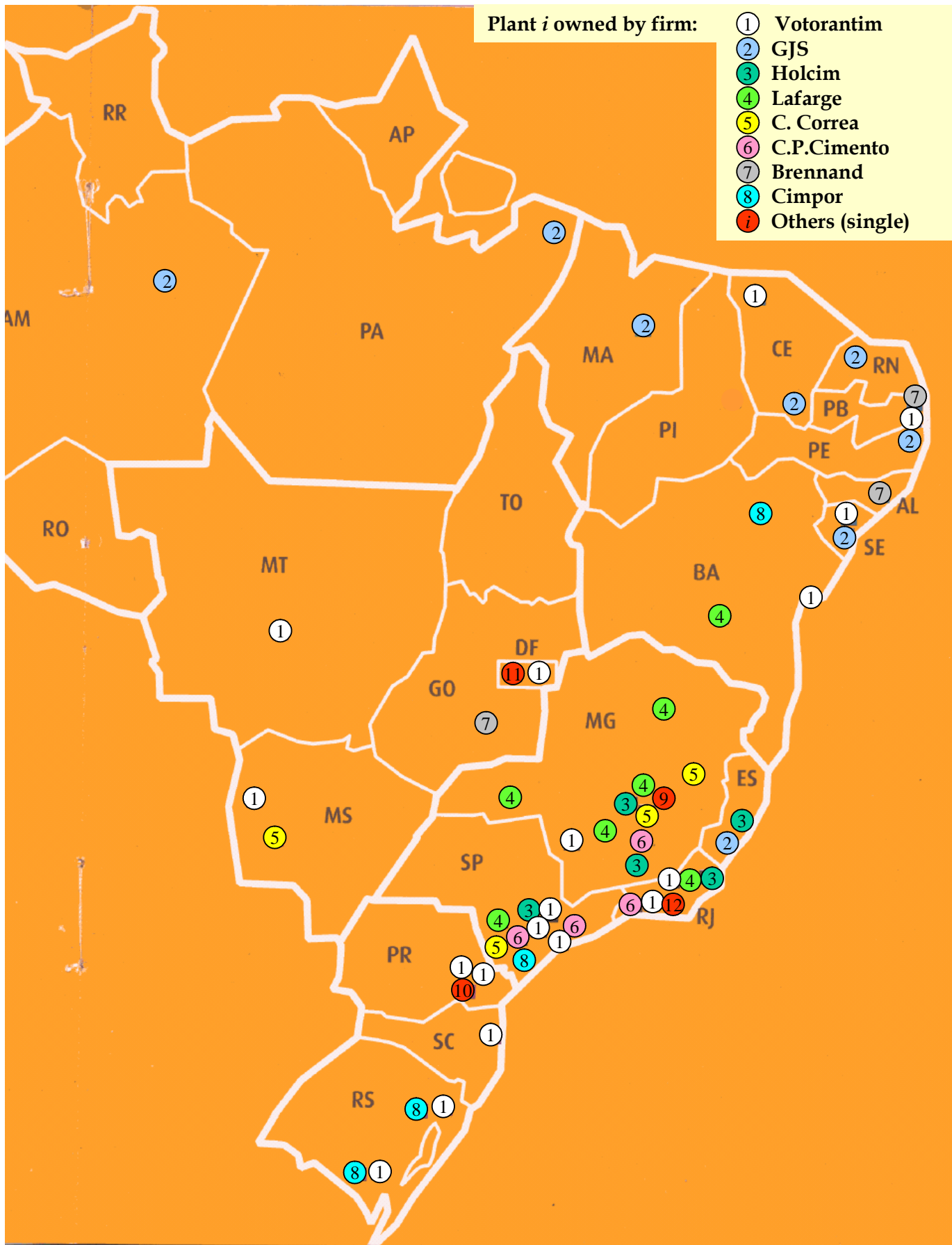
Figure 4: Active plants in 1999

	Mean	Standard Deviation	Maximum	Minimum	Total across 27 states
Cement consumption in state (kt)	1,483	2,324	11,723	55	40,045
Number of (active) cement plants located within state[0]	2.1	2.6	11	0	57
Number of cement firms (producers) shipping to state	5.7	2.8	11	1	12
One-firm concentration index in state[1]	57%	17%	100%	25%	41%
Two-firm concentration index in state[1]	83%	13%	100%	49%	52%
Four-firm concentration index in state[1]	97%	6%	100%	77%	70%
Hirschmann-Herfindahl index in state[1]	4494	1823	10000	1830	2106
% shipments originating from state destined for that state[2]	60%	22%	100%	14%	
% shipments origin. from state destined for that and bordering states[2]	92%	9%	100%	70%	
Value Added (volume decomposition) in Construction Sector[3]	475	726	3,431	9	12,352
Land area (x 1000 square kilometers)[4]	315	370	1,571	6	8,515
Population (m, mid 1999)[4]	6.1	7.3	35.8	0.3	163.9
Population density (/ sq km)	56.9	84.1	339.5	1.2	19.3
Per capita cement consumption in state (kg p.c.)	211	67	353	104	244
Per capita Value Added in Construction Sector[3]	61	26	108	16	75

[0] Of the 57 plants, 7 were grinding-only operations (with clinker being shipped from a nearby plant with integrated facilities)

[1] Based on shipments from producers located anywhere to buyers located in a given state

[2] Applies only to states from which shipments originate (i.e. states where plants are located)

[3] In rescaled constant monetary units

[4] Source: Brazilian Institute for Geography and Statistics (IBGE)

Figure 5: Variation across 27 states of the Brazilian federation, Summary Statistics (time-varying figures refer to 1999)
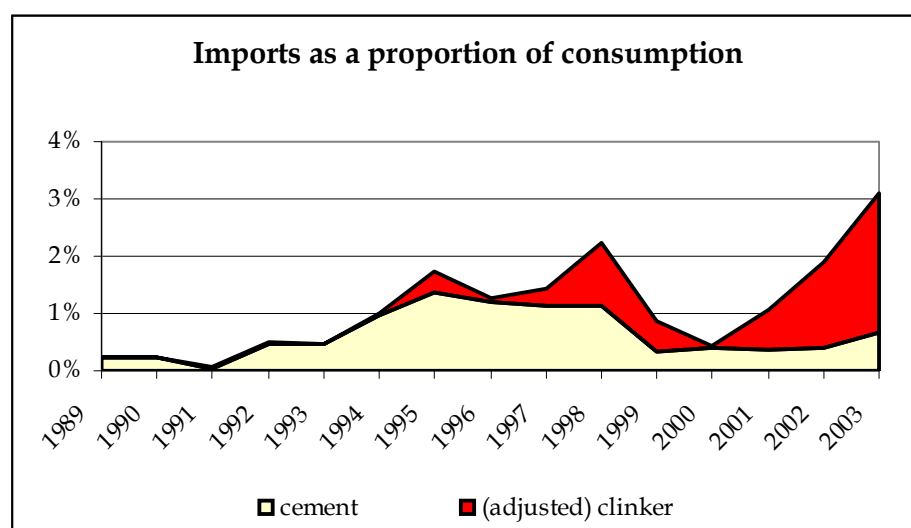

Figure 6: (Official) Imports of cement and clinker as a proportion of domestic consumption. Source: SECEX, MDIC. Clinker quantities are adjusted by the author to reflect usage in the production of cement (assumes 80% of clinker imports used in production of slag cement, with a 40% clinker content).
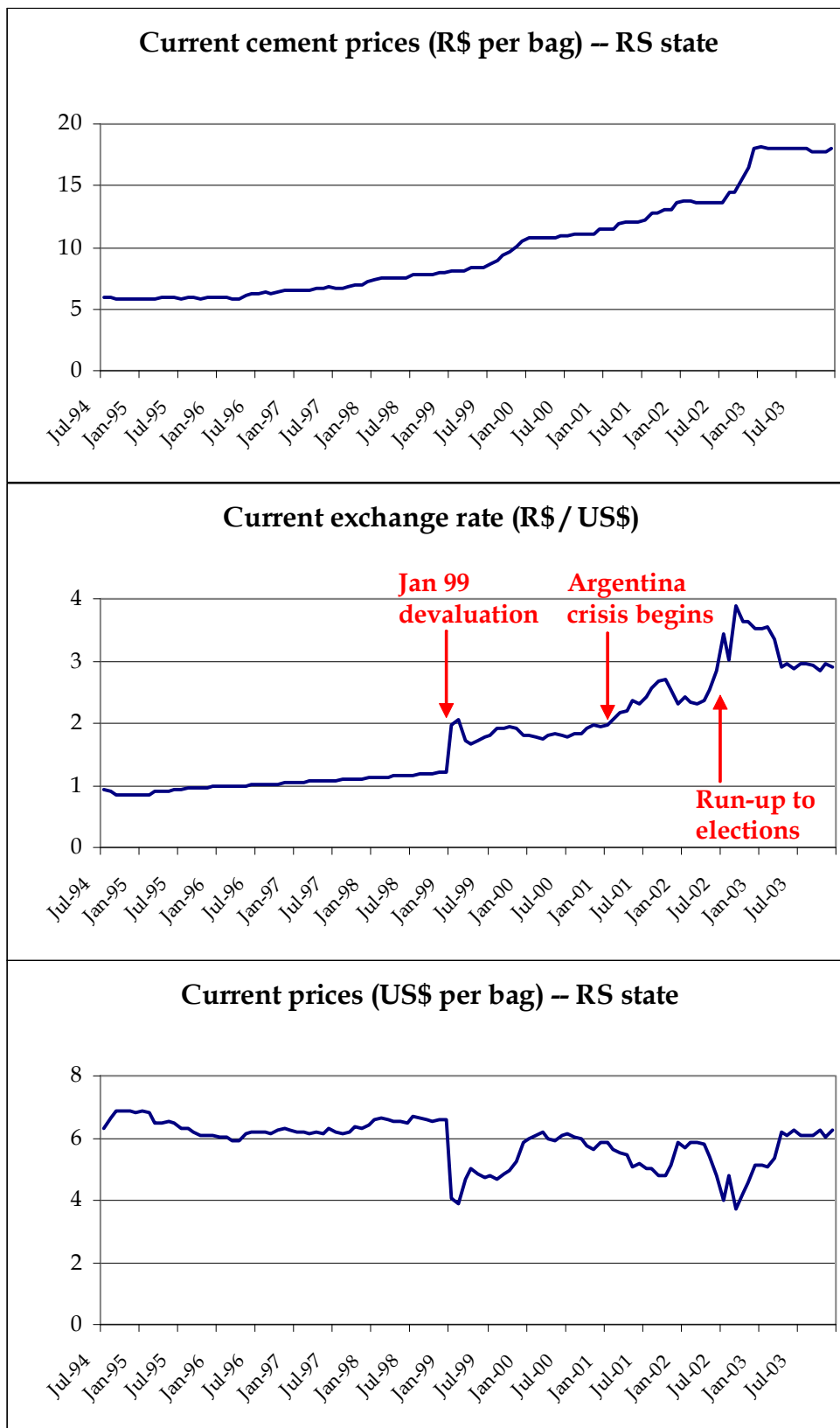
49

Figure 7: Evolution of cement prices in *RS* state since July 1994. In current local currency units (R$) per bag and US$ per bag
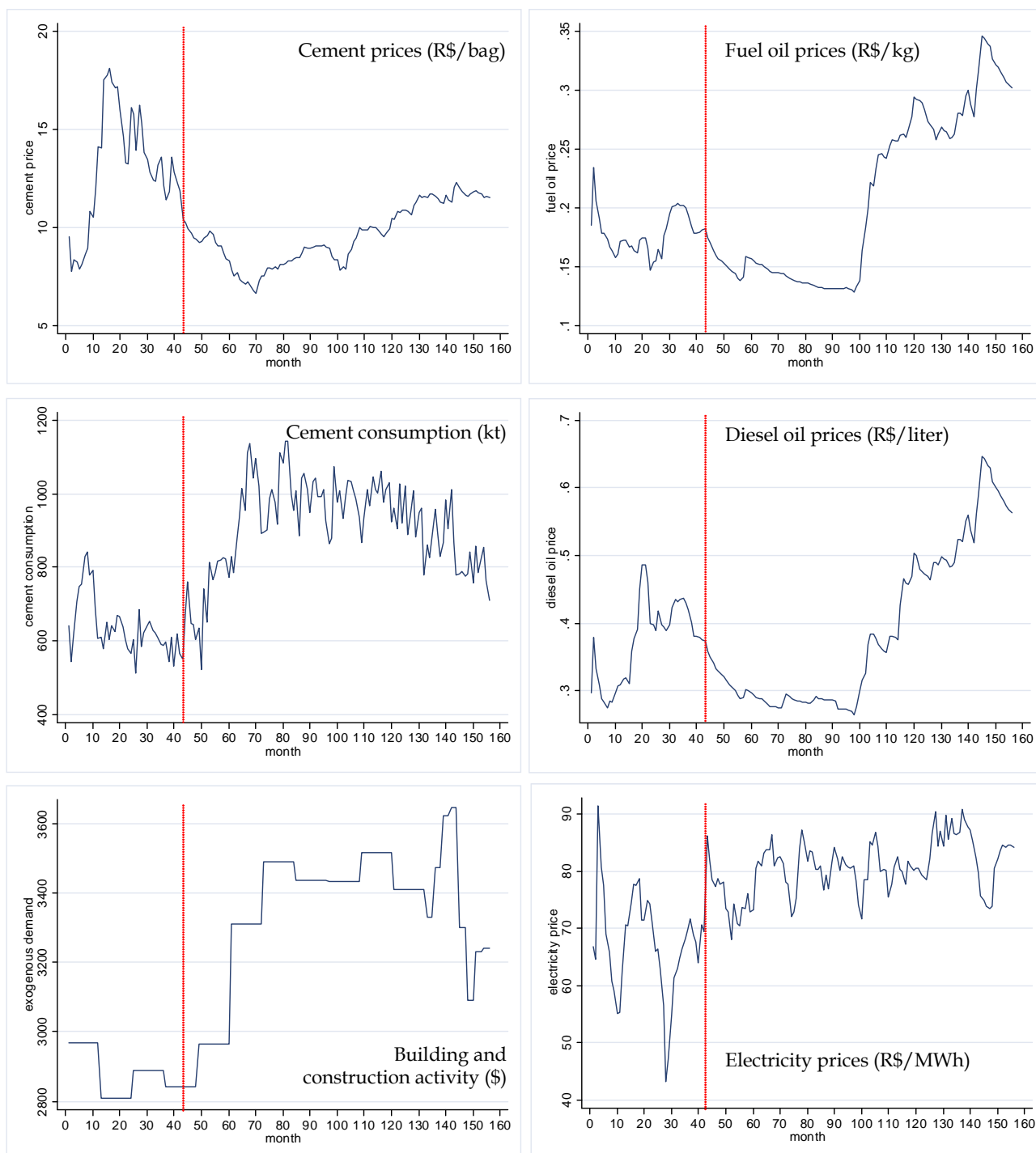
Figure 8: Cement prices, consumption, exogenous demand and factor prices for the state of São Paulo. All prices are in constant December 1999 values. Monthly observations, observation 1 corresponding to January 1991. July 1994, the month in which the stabilization plan was enacted, is marked by the dotted lines.
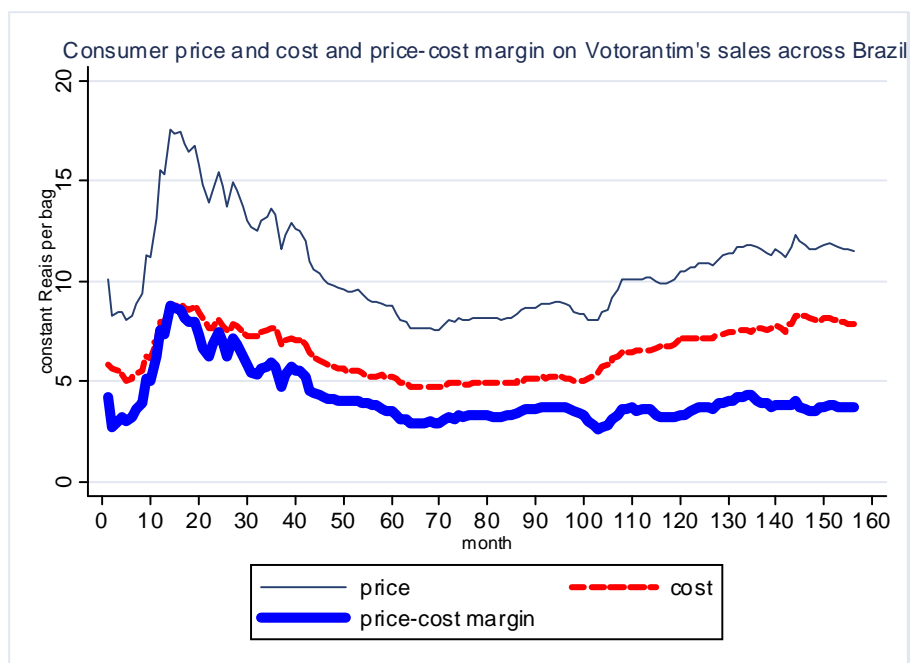
Figure 9: Evolution of consumer prices, marginal costs and price-cost margins on Votorantim's sales. Averaged across all states. In constant Reais per bag (at December 1999 values).

		(I) OLS	(II) IV imports bite	(III) IV imports no bite	(II B) IV subset imports bite	(III B) IV subset imports no bite
No. obs.		156	156	156	156	156
R^2		0.840				
Intercept	coef	2.241 *	2.828 **	2.439 *	0.212	0.729
	s.e.	(1.202)	(1.210)	(1.236)	(1.357)	(1.333)
Exog. demand	coef	0.00159 ***	0.00141 ***	0.00152 ***	0.00225 ***	0.00203 ***
	s.e.	(0.00038)	(0.00039)	(0.00039)	(0.00043)	(0.00042)
Log Price	coef	1.093 **	0.852 *	1.003 *	1.954 ***	1.702 ***
	s.e.	(0.498)	(0.504)	(0.514)	(0.564)	(0.554)
Interaction	coef	-0.000428 ***	-0.000355 **	-0.000396 **	-0.000709 ***	-0.000607 ***
	s.e.	(0.000160)	(0.000163)	(0.000166)	(0.000181)	(0.000176)
Quarterly dummies		Included	Included	Included	Included	Included

Evaluating at the mean of exogenous demand pre-stabilization:
2,883

Intercept	coef	6.825 ***	6.898 ***	6.815 ***	6.699 ***	6.594 ***
	s.e.	(0.143)	(0.144)	(0.145)	(0.155)	(0.167)
Log Price	coef	-0.142 **	-0.171 ***	-0.138 **	-0.091	-0.048
	s.e.	(0.055)	(0.056)	(0.056)	(0.060)	(0.065)

Evaluating at the mean of exogenous demand post-stabilization:
3,338   16% growth versus pre-stabilization phase

Intercept	coef	7.549 ***	7.541 ***	7.507 ***	7.724 ***	7.521 ***
	s.e.	(0.129)	(0.136)	(0.135)	(0.142)	(0.141)
Log Price	coef	-0.337 ***	-0.333 ***	-0.318 ***	-0.414 ***	-0.325 ***
	s.e.	(0.058)	(0.060)	(0.060)	(0.063)	(0.062)
Test of overidentifying restrictions			Fail	Fail	Pass	Pass

Note: Heteroskedasticity and autocorrelation-robust standard errors (Newey-West 1 lag)
    *** Significant(ly different from zero) at the 1% level; ** Significant at the 5% level; * Significant at the 10% level
    Dependent variable is Log Consumption
    Quarterly dummy variables for quarters 1, 2 and 3 are included but estimates are not shown

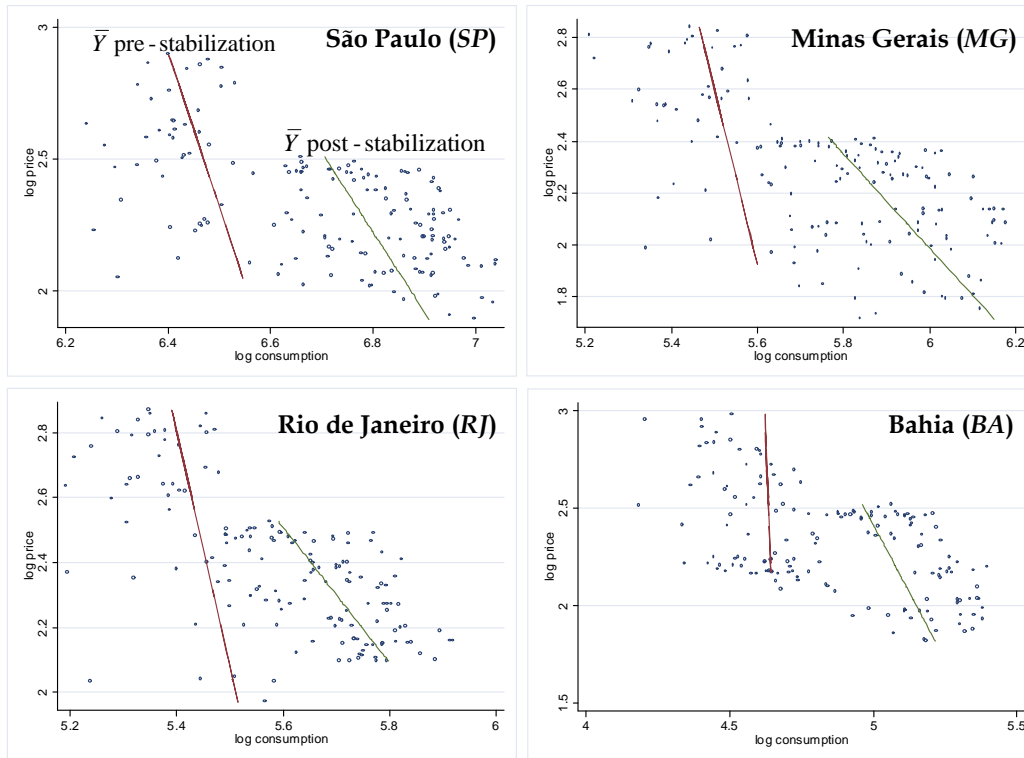Figure 10: Demand estimates for the state of *SP*

Figure 11: Fitted demand curves for the four largest markets. (Log) Price against (Log) Consumption. Evaluated at the respective means of exogenous demand Y for the pre- and post-stabilization phases

		(II) IV-imports bite					
				Log Price: Y evaluated at			
	Cement consumption	Interaction		mean pre-stabilization		mean post-stabilization	
State	in 1999 (kt)	coef	s.e.	coef	s.e.	coef	s.e.
20 SP	11,723	-0.000355	(0.000163) **	-0.171	(0.056) ***	-0.333	(0.060) ***
17 MG	5,090	-0.001067	(0.000235) ***	-0.147	(0.063) **	-0.549	(0.059) ***
19 RJ	3,809	-0.002660	(0.000575) ***	-0.137	(0.059) **	-0.481	(0.057) ***
16 BA	2,461	-0.003048	(0.000815) ***	-0.027	(0.065)	-0.361	(0.079) ***
21 PR	2,321	-0.001015	(0.000647)	-0.137	(0.087)	-0.278	(0.088) ***
23 RS	2,221	-0.001057	(0.000762)	-0.228	(0.037) ***	-0.379	(0.097) ***
22 SC	1,648	-0.003488	(0.002647)	0.020	(0.091)	-0.180	(0.095) *
13 PE	1,225	-0.003389	(0.001675) **	-0.285	(0.093) ***	-0.469	(0.061) ***
10 CE	1,139	-0.005347	(0.001662) ***	-0.142	(0.125)	-0.562	(0.113) ***
18 ES	837	-0.003029	(0.002317)	-0.370	(0.078) ***	-0.480	(0.068) ***
8 MA	765	-0.020114	(0.007056) ***	-0.097	(0.187)	-0.564	(0.126) ***
12 PB	565	-0.036712	(0.007397) ***	-0.123	(0.081)	-0.715	(0.111) ***
11 RN	531	-0.005411	(0.004692)	-0.145	(0.146)	-0.300	(0.078) ***
25 MS	454	0.000899	(0.004419)	-0.431	(0.047) ***	-0.415	(0.071) ***
14 AL	384	0.080309	(0.030990) **	-0.475	(0.127) ***	-0.351	(0.112) ***
9 PI	379	0.015324	(0.012214)	-0.657	(0.272) **	-0.330	(0.103) ***
15 SE	282	0.003937	(0.020794)	-0.145	(0.136)	-0.136	(0.099)

Note: Heteroskedasticity and autocorrelation-robust standard errors (Newey-West 1 lag)

*** Significant(ly different from zero) at the 1% level; ** Significant at the 5% level; * Significant at the 10% level

Figure 12: Demand estimates by state, Summary

| | (I) IV | | (II) IV | |
| | Full sample | | Post-stabilization subsample | |
	coef	bootstrap s.e.	coef	bootstrap s.e.
No. obs.	2652		1938	
Market-specific conduct parameters				
20 *SP*	0.0167	(0.0194)	0.0021	(0.0152)
17 *MG*	0.0194	(0.0127)	-0.0049	(0.0163)
19 *RJ*	0.0079	(0.0206)	-0.0112	(0.0120)
16 *BA*	0.0004	(0.0100)	-0.0268 *	(0.0142)
(Parameters for 13 other markets not shown)				
Factor prices				
Fuel oil (interacted with fuel use)	18.1368 ***	(2.7773)	20.1344 ***	(2.7119)
Coal (interacted with fuel use)	0.0906 ***	(0.0343)	0.0447	(0.0430)
Electricity	0.0343 ***	(0.0125)	0.0494 ***	(0.0169)
Labor	-7.7850 ***	(1.2287)	-3.9898 ***	(1.1363)
Freight (distance interacted with price of diesel oil)	0.0065 **	(0.0028)	0.0066 **	(0.0026)
Other supply-shifters				
Size of sourcing plants	-9.38E-08	(5.84E-07)	4.56E-08	(6.64E-07)
Age of sourcing plants	0.0191	(0.0316)	0.0188	(0.0321)
Price controls (Jan 91 to Oct 91)	-4.4828 ***	(0.8479)		
Intercept (*SP*)	12.1986 ***	(2.6814)	6.9360 **	(2.8436)
(Other market-specific fixed effects included but not shown)				

Note: Heteroskedasticity-robust standard errors with bootstrapping to account for demand estimation
in the first stage. 1000 repetitions, clustered by month (e.g. in (I) a bootstrap sample consists of
156 month draws, and for every month in the bootstrap sample there are 17 markets).
Demand estimates from the first stage of (II) also based on post-stabilization subsample.
*** Significant at the 1% level; ** Significant at the 5% level; * Significant at the 10% level
Dependent variable is the price of cement in units of local currency per bag (at Dec 1999 prices)

Figure 13: Estimation of a static pricing equation, assuming cost is not known. Instrumented with exogenous demand variables.
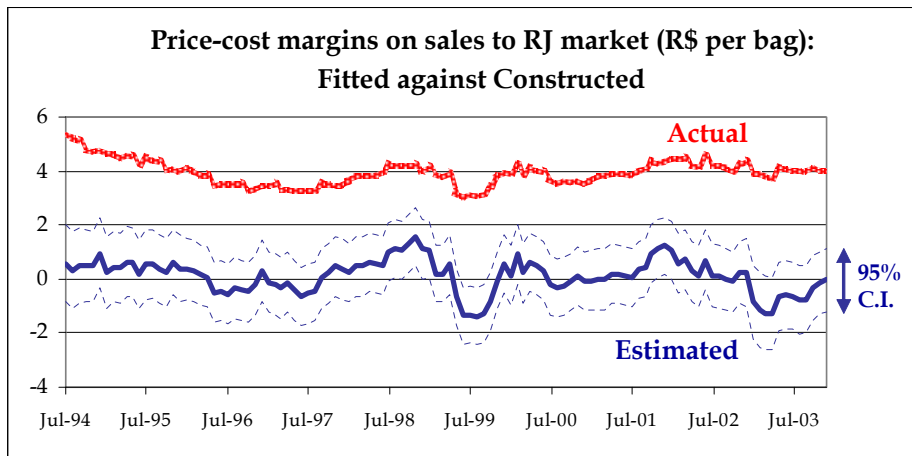
Figure 14: Estimated (average) price-cost margin on sales to RJ market, as estimated by the static pricing equation, against Actual (constructed) price-cost margin. In constant R\$ per bag (December 1999 terms).
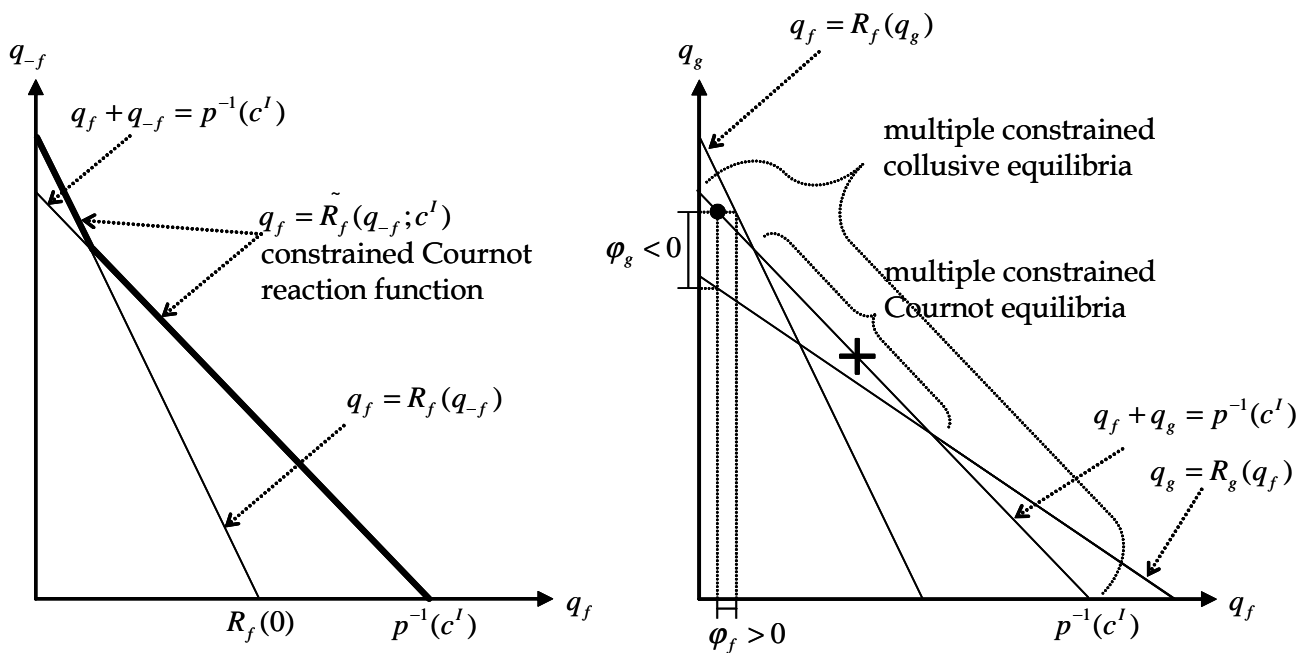

Figure 15: Identifying collusion from Cournot when imports constrain equilibrium prices under both models of conduct. Drawn for a given local market. Left panel: Under the null, Cournot firm f's reaction function, facing domestic rivals and imports. Right panel: Constrained equilibrium outcomes under different behavioral assumptions: Cournot and collusion. The imports constraint binds for both illustrated industry outcomes, marked "$+$" and "$\bullet$". Cournot behavior can be rejected in favor of more collusion only for firm f in outcome "$\bullet$".