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, cost and conduct cannot be identified from the comparative statics of equilibrium. 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 conduct 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. A test of conduct is proposed, adapted to this constrained setting, which suggests that outcomes in the industry are collusive and characterised by market division. The Brazilian cement case also illustrates that incumbents do respond to the *threat* of actions by other agents (such as entrants, consumers and authorities), not only to their *actual* actions.

Keywords: Measurement of market power; Conduct; Limit pricing; Threat of entry; Market division; Cement

JEL classification: L13, L41, L70, F14

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1 Introduction

In the simple version of the static oligopoly model of profit maximisation, firms set prices (or quantities) to solve unconstrained optimisation problems, maximising 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. In a recent study of the airline industry, Goolsbee and Syverson (2005) find that the threat of entry restrains pre-entry prices set by incumbents. 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 optimisation 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 leads to a *downward* bias in the estimated degree of market power. 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 between 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 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 are observed yet the threat of entry (imports) sets an upper limit on prices, equal to the 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 are observed. I take on the role of the researcher who overlooks the latent effect of imports and use the standard methodology to estimate conduct and cost. The supply estimates I obtain are indicative of domestic competition, with price-cost margins centred around zero. Now, the simple technology of the cement industry allows me to observe (construct) 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 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 yields biased supply estimates in the direction of more competition[2].

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 behaviour 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 literature are motivated, for example, by reference to a signalling 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[3]. 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[4]. Indeed, many models in the international economics and macroeconomics literatures take the

[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.

[2] Further, the estimated coefficients on factor prices and other supply-shifters are mostly of the expected sign and significant, which could again mislead a researcher into thinking that the econometric model is appropriately specified. But this owes only to the fact that the estimated coefficients are picking up the expected correlation between cement prices and factor prices.

[3] 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 a downward-biased estimate of market power.

[4] Alternatively, one can view the marginal firm (or imports) as entrants with access to obsolete

world price (plus trade cost) as a domestic limit price, under the implicit *contestability* assumption that the threat of entry is present in the form of opportunistic imports in a well-functioning international trading market, where entry (and exit) are free and there is no entry lag. Of further relevance, macroeconomists have long studied how menu costs and other frictions (e.g. long-term contracts between suppliers 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 "behavioural" 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[5]. 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 realisations of demand such that the cartel's incentive constraint binds and the fully collusive price is not sustainable[6]. Indeed, though his 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.

Given the failure of the standard methodology when market outcomes are constrained, the immediate question is: but what if one *observes marginal cost*? Clearly, a direct comparison of marginal cost to price will provide a test of perfectly competitive behaviour against less competitive models of conduct. But, other than perfect competition, how may one distinguish empirically between two alternative models of firm behaviour when the threat of entry constrains these alternative models to generate the same *aggregate* (market) outcome? In the Brazilian cement industry, for example, where high-cost imports restrain equilibrium prices, one may wish to identify the model of firm behaviour supporting the constrained outcomes. I propose a test of conduct against a standard benchmark, the Cournot oligopoly solution. The measure I develop uses firm-level quantity data to test the hypothesis of Cournot conduct by each firm in each local market against the alternative of "more collusive" behaviour. It is predicated on the notion that no Cournot firm can perceive that marginal revenue (taking rivals' output as given at the margin) exceeds marginal cost, otherwise the firm would optimally expand

technology that have previously exited the industry (with Bertrand pricing between incumbents and entrants).

[5] On the cost of imports, see Harris (1984), who considers the Eastman-Stykolt (1966) hypothesis, whereby providing a domestic industry with a focal point equal to the world price plus tariff, protection may facilitate oligopolistic coordination of the protected firms. On the role of regulated price ceilings in facilitating coordination in credit cards, see Knittel and Stango (2003).

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

output, and this notion holds regardless of whether the imports constraint binds or not in equilibrium[7]. The requirements on the data are large, but the value of the test lies in uncovering *firm-level* behaviour when market outcomes are constrained by the threat of entry and thus the comparative statics of equilibrium are not informative. I illustrate the proposed test of conduct again by reference to the Brazilian cement industry. I find that conduct across firms and local markets is considerably more collusive than the Cournot benchmark, and can be characterised by market division. A story where firm 1 tacitly agrees to give firm 2 the upper hand in market B in exchange for the latter staying away from market A helps to explain the observed shipments. *Importantly, this can be identified despite the threat of imports restraining prices.*

This paper thus makes four contributions. *First*, it demonstrates that the widely-used methodology for inferring conduct leads to the underestimation of market power in industries where equilibrium prices are constrained in ways that are unobserved to the researcher[8]. Potential competition from imports – one of several channels by which such constraints may bind – alone suggests that this result is increasingly relevant for antitrust authorities attempting to measure the competitiveness of conduct in a world where trade barriers are being pulled down. *Second*, I develop a test of conduct for such constrained settings and use it to show that market outcomes in the Brazilian cement industry are indicative of market division. The *third* contribution relates to policy learnings one can derive from the Brazilian cement case. In a developing country such as Brazil, with its huge housing deficit and infrastructure needs, the importance of the cement industry cannot be overstated. Cement is an essential input to construction activity for which there are few substitutes. The industry regularly attracts attention from antitrust authorities, consumer associations and the financial media for its alleged pricing power. A clear policy recommendation is to reduce the transaction costs of imports, thus curbing the substantial market power of domestic producers[9]. *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. I then turn to institutional aspects of the cement industry,

[7]The reverse notion, that a Cournot firm optimally cuts output when its perceived marginal revenue (were imports not to exist) falls short of marginal cost, no longer holds when the price ceiling imposed by imports binds: cutting output in an attempt to raise price above the price ceiling simply invites foreign entry. The test thus endogenously selects from the sample. The key is to understand that in a constrained setting the *first-order-condition off which behaviour is identified is an inequality*, not an equality.

[8]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 18.

[9]By reducing entry costs, one should not expect an increased share of imports – a typical political concern in a country such as Brazil for a "commodity" product such as cement – but rather higher consumer welfare in the form of lower prices. Such a recommendation stands in contrast to recent policy experience, where the cement industry successfully lobbied the government into enacting "antidumping" measures against foreign producers who were attempting to make inroads into Brazil's local markets.

and present the data. Section 4 presents the application[10]. Finally, I conclude, reflecting on the methodological implications of this paper.

2 Theoretical framework

I begin by providing a brief intuitive discussion on why the standard methodology for inferring supply fails when the threat of entry constrains market outcomes. Potential entry is modelled as a competitive fringe of foreign suppliers (imports) to the domestic oligopoly market. I then formally demonstrate the downward bias in the estimated degree of market power. Finally, I specify a test of conduct adapted to such constrained settings.

2.1 The standard methodology under the threat of entry

The static setting considered by Bresnahan (1982, 1989) is depicted in Figure 1 (where marginal cost is flat in quantity, with no loss of generality). 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 c_M, or a high-cost competitive industry, with cost c_C. Then an exogenous shift (rotation) to demand is observed and the equilibrium changes to either E_2^M or E_2^C (changes to E_3^M or remains at $E_1 = E_3^C$) according to the hypothesis of firm behaviour. 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.

Figure 1 about here

Now modify the low-cost monopoly hypothesis by allowing potential entry to constrain prices. The monopolist faces a competitive fringe of imports (labelled I), with perfectly-elastic supply at marginal cost $c_I > c_M$. In general, the equilibrium is given by either of two situations (see Figure 2). If the marginal cost of imports is lower than the monopoly price in the absence of imports (denoted p^M), the price will be equal to the marginal cost of imports, the monopolist will supply the entire domestic market, yet the foreign fringe exerts downward pressure on price. Alternatively, if $c_I > p^M$, 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. Formally, the equilibrium price p is given by

$$p = \begin{cases} c_I & \text{if} \quad p^M \geq c_I \\ p^M & \text{otherwise} \end{cases}$$

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 c_M.

[10] A separate *Supplementary Section*, available from the author, provides a discussion of the sources and treatment of the data used, and all manner of robustness checks regarding the construction of marginal cost and the structural estimation.

Given the assumption that $c_I > c_M$ (imports are high cost), entry is not observed[11]. For $p^M \geq c_I$, 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). Note that around this constrained equilibrium, fluctuations in the marginal cost of imports, say due to fluctuations in the exchange rate, allow one to trace out the market demand curve, since the kinked equilibrium moves up and down along the demand curve.

Figure 2 about here

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 c_M *and* with imports constraining prices at the marginal cost of imports $c_I = c_C$, against that of a high-cost competitive (domestic) industry, with cost c_C (and where the presence of imports becomes irrelevant). Here, prices do not respond differently to demand shocks according to the behavioural 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[12]. Intuitively, the threat of imports constrains the ability of firms with market power to set (market) 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.

Figure 3 about here

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 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.

[11] Clearly, when $p^M \geq c_I$, 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).

[12] Fluctuations in the demand curve can be broken down into rotations around the price intercept and parallel shifts. When marginal cost is flat, identification is possible only from parallel inward shifts of the demand curve, when these shifts are sufficiently large (we are then back to the situation considered in Bresnahan 1982). Identification is not possible for rotations around the price intercept.

Estimation of a static pricing equation In the empirical literature on conduct, the following static pricing equation is typically specified on the supply side[13]:

$$p + \theta q \frac{\partial p(q)}{\partial q} = c \tag{1}$$

where p is price, q is industry output, c is marginal cost and θ is a conduct parameter. One reason why specification (1) 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 internalises the aggregate inframarginal revenue change from a marginal change in output, 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)[14]. Pricing equation (1) can be rearranged to the familiar "elasticity-adjusted Lerner index" (or price-cost mark-up):

$$\theta = -\eta(p)\frac{p-c}{p} \tag{2}$$

where $\eta(p)$ is the market price elasticity of demand. Clearly, such a specification captures the supply decisions depicted in Figure 1, in the absence of imports, under the alternative hypotheses of conduct[15]. Econometrically, (1) may be implemented by including a zero-mean error term ε^s and proceeding to the estimation of

$$p = -\theta q \frac{\partial p(q)}{\partial q} + c + \varepsilon^s \tag{3}$$

where p and q are observed, $\frac{\partial p(q)}{\partial q}$ has previously been consistently estimated, and one wishes to estimate θ and c. Since $q\frac{\partial p(q)}{\partial q}$ is endogenous, one needs to find excluded instruments from (3). In the absence of imports, exogenous demand variables Y will serve as instruments, since they are correlated with the endogenous variable but uncorrelated with the error ε^s (recall Figure 1). Specification (3) is then estimated by IV or GMM and the identifying assumption is

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

[13] Examples include Gollop and Roberts (1979), Roberts (1983), Porter (1983), Suslow (1986), Bresnahan (1987), Brander and Zhang (1990), Parker and Röller (1997), Genesove and Mullin (1998), Wolfram (1999), Nevo (2001) and Slade (2004).

[14] (1) can be specified at the industry or at the firm level. In the latter case, one may include a subscript f for the conduct and cost parameters, to denote the firm. An industry-level pricing equation can be viewed as the average across firms' individual pricing equations (weighted or not by firms' shares), in which case θ has the interpretation of "the average collusiveness of conduct" (Bresnahan 1989). Further, a common alternative to (1) consists of replacing industry output q by firm output q_f in the inframarginal revenue term of firm f's pricing equation. From the first-order condition, the conduct parameter then corresponds to dq/dq_f, which some have interpreted as a "conjectural variation": by this view, upon expanding its output by dq_f, firm f would hold a "conjecture" dq with respect to the resulting aggregate output expansion. For the purpose of my paper, I abstract from the discussion surrounding this interpretation and the related critique that oligopoly theory to date does not underpin a continuum of values for conduct that would support its free estimation (e.g., Reiss and Wolak 2002).

[15] The vertical distance between the (inverse) demand function and the marginal revenue function is equal to $-q\frac{\partial p(q)}{\partial q}$. Under the unconstrained monopoly equilibrium of Figure 1 this distance is always equal to $p-c$. In contrast, in the constrained monopoly equilibrium of Figure 3 this distance will exceed $p-c$.

The problem with the standard methodology arises in the presence of potential imports (entry), since to the extent that imports constrain market outcomes, fluctuations in the demand curve will be correlated with the error in the specified pricing equation. Due to the price ceiling set by imports, the true model – the data generating process – is given by[16]

$$p = \min\left(-\theta q \frac{\partial p(q)}{\partial q} + c + \varepsilon^s, c_I\right) \tag{4}$$

The standard pricing equation which is taken to the data – the estimated model – is, however:

$$p = -\theta q \frac{\partial p(q)}{\partial q} + c + \xi^s \tag{5}$$

where the (mis)specified pricing equation error is denoted ξ^s. The theoretical specification (1) that underlies the estimated model (5) fails to adequately capture the supply decisions (4) of an industry with pricing power facing the threat of high-cost imports. This is summarised in the following proposition.

Proposition 1 *(Non-identification of conduct) 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. In particular, the true degree of market power θ will be underestimated.

Proof. Let $X_1 := -q\frac{\partial p(q)}{\partial q}$. From (4) and (5), the DGP can be rewritten as $p = \min(\theta X_1 + c + \varepsilon^s, c_I)$ and the estimated model is $p = \theta X_1 + c + \xi^s$. We wish to determine $E(Y'\xi^s)$. The error of the estimated model is

$$\begin{aligned} \xi^s &= \varepsilon^s \mathbf{1}[\theta X_1 + c + \varepsilon^s < c_I] + (c_I - \theta X_1 - c)(1 - \mathbf{1}[\theta X_1 + c + \varepsilon^s < c_I]) \\ &= \varepsilon^s \chi + (c_I - \theta X_1 - c)(1 - \chi) \end{aligned} \tag{6}$$

where, for any particular observation, the indicator function $\chi := \mathbf{1}[\varepsilon^s < c_I - \theta X_1 - c] = 1$ when the market equilibrium is unconstrained by the threat of entry (imports have no bite) and $\chi = 0$ when the equilibrium is constrained. (It is clear from (6) that the data generating process is a generalisation of the static model considered by Bresnahan (1982); this static model would correspond to a specific situation where $\chi = 1$ for all market outcomes, i.e. market outcomes are never constrained.) Assuming that the unobserved supply shock ε^s is orthogonal to the excluded exogenous demand variables Y, $E(Y'\varepsilon^s) = 0$, one may write

$$\begin{aligned} E(Y'\xi^s) &= E\left(Y'\varepsilon^s\chi + Y'(c_I - \theta X_1 - c)(1-\chi)\right) \\ &\leq E\left(Y'\varepsilon^s\chi + Y'\varepsilon^s(1-\chi)\right) = E\left(Y'\varepsilon^s\right) = 0 \end{aligned}$$

[16] It is clear from (4) that, ceteris paribus, the likelihood that the imports 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; and (iv) the lower is the marginal cost of imports c_I.

where the inequality follows from the fact that $1-\chi=1$ when $\varepsilon^s \geq c_I - \theta X_1 - c$ (i.e. when the equilibrium is constrained) and $1-\chi=0$ otherwise, along with the assumption that $Y>0$.

Further, let marginal cost be linear in X_2, where X_2 is an $N\times(K-1)$ matrix of observed variables, both exogenous (such as factor prices, including a constant) and endogenous (such as quantity): $c = X_2\beta_2$. 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 $\beta := (\theta, \beta_2)$. The estimated model is then $p = X\beta + \xi^s$. Denote as Z the matrix of instruments, containing the exogenous elements of X_2 and the excluded exogenous demand variables Y, and assume the rank condition for identification holds. The 2SLS estimator is given by

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

Noting that (i) $E(X'Z)$ and $E(Z'Z)$ are positive definite, and (ii) $E(Z'\xi^s)$ contains either 0 or negative elements (since $E(Y'\xi^s)<0$), the application of the law of large numbers to each term along with Slutsky's theorem yields

$$\text{plim } \hat{\beta} < \beta$$

In particular, plim $\hat{\theta} < \theta$. ■

The failure of the orthogonality condition can readily be seen in the linear demand example of Figure 3, as I show in the Appendix.

Proposition 1 has 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[17]. Suppose a researcher fails to realise this price-restraining effect and runs specification (5) on the data, thinking that the data generating process is (3), when it is actually (4)[18]. 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. The estimated conduct parameter $\hat{\theta}$ will lie below the "true" value $\theta = \eta(p)\frac{p-c}{p}$, as defined in (2), underestimating the degree of market power. Intuitively, since the response of prices to demand shocks is constrained, the coefficient on $-q\frac{\partial p(q)}{\partial q}$ will be biased toward zero. In general, unless the researcher is somehow sure

[17] In the notation of the proof, this corresponds to $\Pr(\chi=0)=\Pr(\varepsilon^s \geq c_I - \theta X_1 - c) > 0$ in the available sample. In contrast, the static model considered in Bresnahan (1982) is a special case of the DGP (4), where no outcomes are constrained: $\Pr(\chi=0)=0$.

[18] 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", he could implement the standard methodology using the unconstrained outcomes only (i.e. where $\chi=1$), or in principle he could use switching regression techniques (assuming, of course, that *not all observations are constrained*). Typically (i) the constraining effect of potential entry may be overlooked since entry is not observed in equilibrium; (ii) the level at which the constraint binds fluctuates and is unobserved (e.g. in the example of footnote 3, 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) it appears that a necessary condition for identification 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.

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 suggests that to the extent that market outcomes in the sugar industry were constrained by the threat of entry, this would lead to a downward bias in the estimated degree of market power[19].

2.2 Testing Cournot against "more collusive" conduct

In a constrained equilibrium, it may be possible to empirically distinguish between alternative models of conduct which, owing to the binding constraint, give rise to the same aggregate outcomes. I propose a measure that uses each firm's supply decision toward each local market to test the hypothesis of (constrained) Cournot conduct against the alternative of "more collusive" behaviour[20]. While the test requires the observation of domestic cost and is only sufficient to reject the hypothesis of Cournot, its value resides in uncovering firm-level behaviour in a constrained setting.

Figure 4 establishes the Cournot benchmark (in any given local market)[21]. 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 constraint posed by the threat of entry on firm f's output, given by $q_f + q_{-f} \geq p^{-1}(c_I)$. (To see this, note that the price ceiling set by the perfectly-elastic supply of imports is equivalent to an output floor.) Thus, firm f's best response to the output of its rivals and to the threat of imports corresponds to the outer envelope to its reaction function in the absence of imports and the boundary to the imports constraint; denote this "constrained" reaction function as

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

Drawn as a thick curve in the left panel, for $R_f(0) < p^{-1}(c_I)$ (such that $q_f = R_f(q_{-f})$ crosses $q_f + q_{-f} = p^{-1}(c_I)$), 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:

$$p(q) + \frac{p(q)}{\eta(q)} \frac{q_f}{q} = c_f \tag{7}$$

[19] 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 $\eta(p)$ of around -1.05 for most part of the year) is surprisingly low, the authors attributing this to the threat of entry.

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

[21] For a model with a similar flavour where a Cournot oligopoly may deter entry by producing the limit output, see Gilbert and Vives (1986) (I thank Xavier Vives for pointing this out to me). As in the empirical literature on conduct, the Cournot assumption serves as a benchmark (e.g. Parker and Röller 1997). See the *Supplementary Section* for a discussion of the suitability of an alternative benchmark model of the Hotelling-Salop type, where pricing by firms across local markets is highly restrictive.

where, as before, $\eta(q)$ is the market price elasticity of demand and $q = q_f + q_{-f}$. Now, for lower q_{-f} such that $\tilde{R}_f(q_{-f}; c_I) + q_{-f} = p^{-1}(c_I)$, the imports constraint binds and firm f's optimal reply *exceeds* the quantity that it would set in the absence of imports. Here, firm f's (perceived) marginal revenue falls short of marginal cost:

$$p(q) + \frac{p(q)}{\eta(q)} \frac{q_f}{q} < c_f \tag{8}$$

Conditions (7) and (8) combine to prove the following proposition:

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

$$p(q) + \frac{p(q)}{\eta(q)} \frac{q_f}{q} \leq c_f \tag{9}$$

This condition holds as a strict inequality when the "imports 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.[22]

The rival firms' constrained joint reaction function, $q_{-f} = \tilde{R}_{-f}(q_f; c_I)$, is similarly derived. The set of equilibrium outcomes is then the intersection of $q_f = \tilde{R}_f(q_{-f}; c_I)$ and $q_{-f} = \tilde{R}_{-f}(q_f; c_I)$. For a low enough cost of imports (i.e. an imports boundary sufficiently far from the origin), there are multiple equilibria and imports restrain prices at the Cournot equilibrium[23].

Figure 4 about here

In a constrained setting, how can one then use observed marginal cost in FOC (9) to test Cournot against the alternative of "more collusive" behaviour? The left panel of Figure 5, depicting domestic duopolists f and g, shows that this may not always be possible. 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. Suppose that one observes the constrained outcome marked with a "+". Because such an outcome is consistent with both Cournot and more collusive behaviour, it is not possible to tell whether the duopolists behave in Cournot fashion (in which case aggregate equilibrium output in the absence of imports would equal $q^C := q_f^C + q_g^C$) or whether firm behaviour is more collusive than the Cournot benchmark (in which case the aggregate equilibrium output in the absence of imports would be lower than q^C). However, it may be possible to observationally distinguish collusion from Cournot if firm-level outputs are "sufficiently" asymmetric[24]. This is

[22] Condition (9) 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.

[23] In the absence of imports, the unique Cournot equilibrium outcome (q_f^C, q_{-f}^C) is defined implicitly by $q_f^C = R_f(R_{-f}(q_f^C))$ and $q_{-f}^C = R_{-f}(R_f(q_{-f}^C))$. Formally, imports have bite under Cournot conduct if $p(q^C) \geq c_I$, where $q^C := q_f^C + q_{-f}^C$, or equivalently when $q^C \leq p^{-1}(c_I)$.

[24] In a non-cooperative framework where side payments are not allowed, with heterogeneous firms meeting in multiple markets, market division can be supported in equilibrium. By pooling incentive constraints across the different markets, a firm's share in those markets where it enjoys a low cost (e.g. those local markets in proximity to its plants, in a spatial model with multiplant firms) may be increased at the expense of its share in markets where it has a high cost (e.g. markets further away from its plants, but closer to the plants of rival firms). See Bernheim and Whinston (1990) and, for an application to the Brazilian cement industry, see the *Supplementary Section.*

illustrated in the right panel of Figure 5. From the observed equilibrium outcome, firm g's behaviour is still consistent with both Cournot and more collusive conduct. Under the Cournot hypothesis, firm g does not cut output in an attempt to raise price above c_I, as an unconstrained Cournot firm would do, since this would only open the door to imports. In contrast, firm f's behaviour is *not* consistent with Cournot. Firm f is restricting output as compared to the output decision of a constrained Cournot firm. The point is to recognise that for a Cournot firm, the general (i.e. allowing for the constraining effect of entry) pricing condition (9) of Proposition 2 has to hold. That is, for no Cournot firm can (perceived) marginal revenue *exceed* marginal cost, otherwise the firm would optimally expand output, and this holds irrespective of whether the imports constraint binds or not (since the latter places a *lower* bound on aggregate domestic output). This translates into the following test. Rewrite (9) as an equality:

$$p(q) + \frac{p(q)}{\eta(q)} \frac{q_f}{q} = \varphi_f + c_f \tag{10}$$

Proposition 3 *(Sufficient statistic to reject Cournot behaviour) Under the null of Cournot behaviour, $\varphi_f \leq 0$. When the imports constraint binds, $\varphi_f < 0$ is consistent with Cournot behaviour. The finding that $\varphi_f > 0$ allows one to reject the hypothesis that firm f is behaving in Cournot fashion, in favour of more collusive behaviour, regardless of whether the imports constraint binds or not.*

Figure 5 about here

3 Industry and data[25]

3.1 The Brazilian cement industry in the 1990s

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 6, in 1999 57 active plants were scattered across a geographic area slightly smaller than that of the US[26]. This spatial distribution is not even, however, as consumer markets and thus plants are concentrated along the coastal states, in particular the relatively wealthy and populated states in the Southeast and South regions of the country[27]. States to the northwest of the centre of the country are sparsely populated and are largely covered with jungle.

Figures 6 and 7 about here

[25] This section covers key features of the Brazilian cement industry and of the data over the relevant period. General technological and demand-side considerations of the industry, as well as a detailed account of the sources and treatment of the data, are provided in the *Supplementary Section*.

[26] 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).

[27] 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).

In 1999, as also depicted in Figure 6, these 57 plants were owned by 12 firms. The two largest firms, Votorantim and Grupo João Santos, respectively with nationwide shipment shares of 41% and 12% in 1999, were both of Brazilian origin. The subsidiaries of the large multinational firms Holcim and Lafarge followed, with shipment shares of 9% and 8% respectively. As Figure 7 indicates, this national picture hides a lot of variation at the local, statewide level.

The 1990s saw two distinct periods in the history of the Brazilian cement industry. 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 stabilisation 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[28]. Thanks to the industry's practice of investing in idle capacity, firms were able to meet rising demand[29]. The post-stabilisation phase of the 1990s also saw a flurry of acquisition activity in the cement industry, with the expansion of incumbents and the entry of foreign firms which did not previously own assets in Brazil. Compared to the 12 firms that ran operations by 1999, the industry had consisted of 19 producers in 1991.

Given the short shelf life of cement, firms produce for immediate consumption. Stocks at producers amount to approximately one week of sales, with roughly another week of sales being stocked down the trade. Around 90% of shipments from producer plants to buyers in consumer markets is carried out by road – as opposed to rail or water. In line with other developing countries, and in stark 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[30]) constitutes a small share of domestic consumption. As shown in Figure 8, in the period 1989 to 2003, this share has amounted to at most 2-3% of consumption across Brazil, though the trend appears to be rising since the trade liberalising reforms of the early 1990s (and despite a dip in 1999 and 2000 following the devaluation of the local currency – see below). 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

[28] The *Real* economic stabilisation plan quickly brought inflation under control. The sharp slowdown in inflation, through the reduction in "inflation tax", represented a reduction in the transfers from the private sector to the government. In particular, the large mass of consumers among the lower-income groups 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. See Salvo (2004) for further details.

[29] As I present shortly, imports remained contained and prices actually fell in this period. Capacity utilisation typically hovers around an average 65%. See the *Supplementary Section* for further details, including a discussion of a strategic role for capacity in this industry.

[30] Clinker is the main component of ordinary cement (see the *Supplementary Section* for details).

markets is actually as high as 30%[31]. The presence of imports in Brazil thus pales in comparison to the US, despite most of its markets being located along (or in proximity to) an extensive Atlantic coastline.

Figure 8 about here

As this study finds, however, the limited penetration of imports hides their role in restraining domestic prices. The trade liberalising reforms of the early 1990s, coupled with the appreciation of the local currency in the four years following stabilisation in mid 1994, opened the door to the threat posed by the entry of imports. To provide a flavour, Figure 9 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 (the *real*, R$) and in a currency of foreign trade, proxied by the US dollar. Domestic cement prices in local currency are highly correlated with the price of the US dollar in local currency (i.e. the exchange rate), to the extent that despite the occurrence of large variations in the exchange rate during the period, the domestic cement price converted into US dollars is quite steady since 1995[32] [33]. Also consistent with the imports-constraining story, in the cross-section of local markets cement prices are increasing in the market's distance from the coast. But I do not take either of these observations – the correlation between cement prices and the exchange rate, or the cement price increasing in distance – as *prima facie* evidence that potential imports restrain prices. These observations could also be explained by alternative stories, such as factor prices being set in hard currency on the world market (i.e. oil), 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 supported by interview evidence, will be the key element in support of my claim.

[31] Despite the bulkiness of cement relative to its price, the development of specialised seaborne handling and transportation equipment from the 1970s enabled imports to make their presence felt in coastal markets. Dumez and Jeunemaître (2000) provide a historical account of the rise of international trade in cement. On the other hand, in both the US and Brazil, exports account for less than 1% of domestic production (though in Brazil the current trend is upwards).

[32] Until January 1999, Brazil had an exchange rate fixed by the government. The local currency was then floated, in the midst of the "Brazil currency crisis", depreciating by 70% against the US dollar in one month, but later partially receding. Other periods of abnormal 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 outcome of 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, as defined in Section 2). The observation 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 devaluations in the local currency (i.e. in 1999, in 2001 and in 2002) suggests that raising domestic prices in local currency is not friction free (perhaps the industry is wary of attracting negative publicity).

[33] To provide an example, an equity analyst wrote that "(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" (Zaghen 1997; p. 24). (The author refers to the exporter's price for delivery "at the coast", excluding cost upon arrival in Brazil, such as inland freight, sales taxes and resellers' markups.) Further evidence that domestic producers were threatened by imports is their successful lobbying of government in passing antidumping measures – namely a 23% import tariff – against Venezuelan and Mexican cement producers in the late 1990s who were starting to make inroads into local markets particularly in the north and northeast of the country.

Figure 9 about here

3.2 Data available: Plant-to-market cement flows and the construction of marginal cost

On the demand side, I observe monthly cement consumption and consumer prices (i.e. prices set by resellers) across the 27 states in the period 1991 to 2003. I take each state to represent a local market. As demand shifters, I observe alternative series of economic activity, either in the construction and building sector or aggregated across sectors of the economy, which I use as proxies for the exogenous demand for cement[34].

The key ingredient on the supply side is the observed breakdown of shipments from each plant to each of the local markets (states), enabling me to map the flow of cement from the plant to the consumer. In addition to plant ownership, I observe plant characteristics – e.g. capacity, number of kilns, type of fuel usage, proportion of shipments in bags as opposed to bulk – and local factor prices, such as fuel oil, coal, electricity and wages. I do not observe freight prices paid by cement producers but I approximate these by using data on freight prices for agricultural goods collected over the period 1997 to 2003 for thousands of different routes across Brazil. The transportation of goods such as soyabean and maize are reportedly close substitutes in the supply of cement freight (Soares and Caixeta Filho 1996).

Considering that the technology of cement production is of the fixed coefficients type, I use engineering estimates, factor prices and plant characteristics to directly calculate the marginal cost of each plant in serving each market[35]. In view of the technology and my understanding of the industry, I model plant marginal cost as flat in quantity up to capacity. Notice that I do not observe producer prices, only consumer prices. However, I back out producer prices assuming competition at the retail (reseller) level and taking into account the high proportional sales taxes. The assumption of competition among resellers follows from several field interviews, including interviews with producers' sales representatives and with resellers. (I also check the robustness of this assumption by, for example, comparing *observed* producer prices that I was able to obtain from a subset of producers to the backed-out producer prices.) This study thus considers 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.

Figure 10 depicts cement prices (in units of local currency for the standard 50 kg bag, at a constant December 1999 level[36]), cement consumption and exogenous demand (activity in the construction sector) from January 1991 through December 2003 for the

[34] This follows from the fact that cement is an input to construction and yet accounts for a small share of construction budgets (see the *Supplementary Section*). Taking such construction activity to exogenously move the demand curve for cement is a typical assumption: see, for example, Syverson (2004) who uses construction sector employment as an exogenous measure of demand in ready-mix concrete (an industry located downstream to cement).

[35] As I argue in the *Supplementary Section*, my calculations of marginal cost actually overstate the true marginal costs. But when I turn to the testing of conduct, such a bias reinforces the results.

[36] This is done using an economy-wide General Price Index (GPI). Owing to the high levels of inflation prevailing in the first 42 months (out of 156) that I consider, particular attention has been paid to the conversion of current cement prices to constant prices – see the *Supplementary Section*. Factor prices are similarly converted. In contrast, Figure 9 presents current prices (albeit for another state).

largest market, the state of São Paulo (SP)[37]. The month in which the stabilisation plan was enacted, July 1994, corresponds to observation (month) 43 in the graphs (marked by dotted lines). 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-stabilisation period they gradually declined back to R$ 7 by late 1996, gradually rising thereafter. The sharp increase in consumption following stabilisation, from a level of 600 mt per month to 1000 mt per month within two years, pulled by a 20% jump in the level of construction activity, is evident from the graphs. Some factor prices are also portrayed. Of note, there is high correlation in the post-stabilisation phase between the price of cement and the prices of fuel oil and diesel oil[38]. 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 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).

Figure 10 about here

A glance at price-cost margins With respect to firm profitability, Figure 11 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 local currency units per bag[39]. 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, including freight, sales taxes and the reseller's cost. The picture is similar across firms. *In sum, the industry wields considerable market power, despite the threat of imports.* 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)[40].

Figure 11 about here

[37] In addition to accounting for around one-third of the nation's cement consumption, the case for the state of São Paulo is broadly representative for Brazil as a whole.

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

[39] The *Supplementary Section* breaks this aggregate figure for Votorantim down into *state-specific* prices, costs and margins over time, for each of the 25 states where the firm is present.

[40] Two robustness tests of the calculated marginal costs and the resulting price-cost margins are conducted in the *Supplementary Section*. 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.

4 Inferring demand and conduct in the Brazilian cement industry

4.1 A "road map"

I begin in Section 4.2 by estimating demand in each local market (state). The market price elasticities of demand are estimated to be very low, of the order of -0.5, including local markets where the one-firm concentration ratio is as high as 80%. Two main possibilities arise to rationalise why an industry facing such inelastic demand does not cut output to raise prices to a point where demand is more elastic: (i) there is weak pricing power (e.g. competition or low concentration), or (ii) some dynamic story is appropriate, such as the threat of entry (imports) restraining pre-entry prices[41]. 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. The second explanation is the one I accept. 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 elastic[42]. Attempts by the domestic industry, already enjoying a large price-cost margin, to raise prices above this price ceiling would only invite foreign entry[43]. This story is further supported by a wealth of anecdotal and interview-based evidence. It is also consistent, as argued in Section 3.1, with the high correlation observed between cement prices and the exchange rate, the latter having varied considerably over the time period.

In this constrained setting, Section 4.3 then illustrates the poor performance of the standard methodology for inferring conduct and cost. 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. The negative bias in the estimated price-cost margins is severe in light of the high observed price-cost margins.

Finally, Section 4.4 delves deeper into the pattern of conduct in the industry, taking into account the constraint posed by imports on industry outcomes. I use the test of

[41] A third possibility hinges on a very special class of models of spatial competition, where a firm is restricted to set only a "mill" price, with delivered prices to consumers who are distributed over space being equal to the sum of this mill price and the transportation cost. See the *Supplementary Section.*

[42] 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.

[43] Other studies of cement have found low market price elasticities in equilibrium. For example, Röller and Steen (2002) find an average fitted 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 have also been restraining the prices set by the domestic oligopolies. The explanation commonly advanced behind such inelastic demand is that cement accounts for a low share of construction budgets and 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.

Proposition 3, which is based on direct measures of costs and the observed firm-level quantity data (i.e. the flow of cement from plants to local markets), to show that conduct is considerably more collusive than the Cournot benchmark.

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 $i = 1, ..., I$.[44] Let $i = 0$ index the aggregate fringe of foreign suppliers. The flow of cement for consumption can be summarised 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 in each market l can then be written:

$$q_{lt} = D(p_{lt}, Y_{lt}, \alpha_l, \epsilon_{lt}^d) \tag{11}$$

where p_{lt} is the price of cement to the consumer, Y_{lt} are exogenous variables shifting demand (e.g. output in the construction and building sector), α_l are market-specific parameters to be estimated and ϵ_{lt}^d is an econometric error term. (Demand function $D(.)$ is the inverse of the inverse demand function $p(.)$ considered in the theoretical framework, i.e. $D(.) = p^{-1}(.)$.)

Estimation of (11) 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 behaviour of domestic firms[45]. There are therefore two situations to consider.

Identification 1: Imports restrain domestic prices at the industry equilibrium In practice, due to the presence of frictions, cement prices will not be exactly equal to 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 the marginal cost of imports allow one to trace out the demand curve (assuming c_I does not rise to the extent where imports no longer have bite). The marginal cost of imports 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 be used as instruments for prices in the estimation of (11) (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 prices not determined by the marginal cost of imports c_I – is orthogonal to the unobserved demand shocks across local markets, prices can be treated as predetermined and (11) can be estimated by OLS[46].

[44] Not all plants are active (in the sense that cement is shipped from them) in each time period t; in a given time period, some plants may be yet to enter (or reenter), while others may have exited.

[45] For example, if $q^M \leq p^{-1}(c_I) < q^C$, imports have bite under full collusion but not under Cournot.

[46] The model I have in mind when imports restrain domestic prices is as follows. Cement prices p are

Identification 2: Imports do *not* restrain domestic prices 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 (flat) marginal cost.

Demand specification I fit alternative parametric specifications for the market-level demand function (11), such as the loglinear form[47]:

$$\log q_{lt} = \alpha_l^1 + \alpha_l^2 Y_{lt} + \alpha_l^3 \log p_{lt} + \alpha_l^4 Y_{lt} \log p_{lt} + \epsilon_{lt}^d \tag{12}$$

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, (12) 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 12 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[48]. The (average) market price elasticity of demand during the pre-stabilisation phase amounts to (an inelastic) -0.17, rising to -0.33 during the post-stabilisation phase[49]. Thus, as prices in the economy stabilise and an average 16% exogenous increase in the demand

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 $E(V^1\epsilon^d) = 0$, where ϵ^d captures the unobserved demand shocks in (11), V^1 (e.g. the exchange rate) can be used to instrument for prices in the estimation of (11). In addition, if $E(\zeta\epsilon^d) = 0$ (and of course $E(V^2\epsilon^d) = 0$ as well), demand equation (11) can be estimated consistently by OLS.

[47] The following comments are of relevance to the results that I present below: (i) though I illustrate using estimates for the loglinear specification, results are robust to the choice of alternative functional forms (see the *Supplementary Section*); (ii) I provide estimates for the state of *SP* in greater detail, as this is the largest market, 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) given the quarterly seasonality of sales, three quarterly dummies – not shown in (12) – are included; (v) standard errors are heteroskedasticity and autocorrelation-robust (1-lag Newey-West errors). I discuss specification tests (heteroskedasticity, serial correlation, overidentifying restrictions, endogeneity) in the *Supplementary Section*. Of note, to check the extent to which overidentification may be driving efficiency at the expense of consistency, I reestimate "base" regressions using subsets of the sets of instruments, to obtain similar elasticities (estimates for these are shown in Figure 12 as regressions (II B) and (III B)).

[48] Estimation of specification (11) *excluding* the interaction term renders a significantly positive coefficient on the level of exogenous demand, as expected (results are not shown).

[49] Each fitted equation is evaluated at two different values for the exogenous demand variable: at the mean for the pre-stabilisation (high inflation) phase, $\bar{Y}_{SP,pre} = 2883$, from January 1991 through June 1994 (42 observations), and at the mean for the post-stabilisation (low inflation) phase, $\bar{Y}_{SP,post} = 3338$, from July 1994 through December 2003 (114 observations). The reported elasticities (-0.17 and -0.33) can then be read from the coefficients on log price in the bottom half of column (II), computed as $\hat{\alpha}_{SP}^3 + \hat{\alpha}_{SP}^4 \bar{Y}_{SP,pre}$ and $\hat{\alpha}_{SP}^3 + \hat{\alpha}_{SP}^4 \bar{Y}_{SP,post}$. Clearly, a formal test that the price elasticity has increased is equivalent to verifying that the coefficient on the interaction term is significantly negative. This is so:

for cement occurs, the price elasticity doubles from around -0.2 to around -0.4. Figure 13 plots the fitted demand curve evaluated at the means for exogenous demand in the pre- and post-stabilisation 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 stabilisation 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.

Figures 12 and 13 about here

Results by state Figure 14 summarises results across 17 states, from regression (II), again using the full sample (pre- and post-stabilisation phases). (The 10 remaining northwestern states – sparsely populated and largely covered in jungle – are dropped from the analysis owing to measurement error[50].) 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-stabilisation phase, the price elasticity of demand is negative for all 17 states, and significant at the 1% level in 15 states. Post-stabilisation 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. Of note, 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-stabilisation phase is negative in 16 out of 17 states, 9 of which are significant at the 10% level or higher. The mean pre-stabilisation price elasticity across states is lower: -0.22.[51]

In sum, regardless of the type of price instruments employed (or using prices themselves, under OLS), the choice of which depends on whether one accepts that imports restrain domestic prices at the industry equilibrium, I estimate very low market price elasticities of demand, of the order of -0.5.[52]

Figure 14 about here

the p-value for this (one-tailed) test is 1.5%. Importantly, to check the robustness of the low elasticity I repeat regressions (II) and (III) using only the 114 observations from the post-stabilisation subsample (July 1994 on). This confirms a low elasticity of -0.4 (results are not shown).

[50]To provide a point of comparison, as I explain in Salvo (2004), the global market research firm ACNielsen, well-known to marketing professionals in consumer goods industries, and with over 30 years of experience in Brazil, does not audit these states given their unusual geo-demographic characteristics. The 10 states I drop together account for 60% of Brazil's land mass but only 11% of its cement consumption.

[51]The price elasticity increases upon stabilisation in 13 out of the 17 states, in 8 of which at the 5% level of significance. (This follows from the negative interaction coefficients and the fact that exogenous demand rises in *every* state concurrent with stabilisation.) See the *Supplementary Section* for one possible explanation behind this.

[52]The *Supplementary Section* considers further robustness tests, such as fitting alternative functional forms, using fixed-effects IV panel data estimation and reversing the dependent variable.

4.3 The standard methodology: Inconsistent supply estimates

I proceed to estimating a pricing equation such as (3), as is standard in the empirical literature on market power[53]. In order to complete the specification of the structural econometric model, write the market-level demand function (11) in inverse form ($D^{-1}(.) = p(.)$) as

$$p_{lt} = p(q_{lt}, Y_{lt}, \alpha_l, \epsilon_{lt}^d)$$

and define plant i's costs as

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

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 Z_{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. 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 maximise 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:

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

where θ_{flt} is a firm-level conduct parameter which encompasses alternative models of conduct, as discussed in Section 2.1. For $q_{ilt} > 0$ (i.e. an interior solution) and specifying an additive econometric pricing error, one may implement this pricing equation as

$$p_{lt} = -\theta_{flt} \frac{p_{lt}}{\eta_{lt}} + c(q_{it}, \mathbf{q}_{it}, W_{it}, Z_{it}, \beta) + \varepsilon_{ilt}^s \tag{13}$$

In what follows, I present estimation results corresponding to a market-level counterpart to the plant-level pricing equation (13)[54]. In view of the fixed-coefficient technology

[53] Since I now ignore that the true model is (4) and that the estimated model is thus (5) with $E(Y'\xi^s) < 0$, here I refer to the pricing equation error as ε^s and not ξ^s, as a researcher overlooking the price-constraining effect of imports would do, thinking that he is specifying (3) with $E(Y'\varepsilon^s) = 0$.

[54] As mentioned in footnote 14, 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.

of production, I specify average market marginal cost c as being linear in average market factor prices W_{lt} (namely fuel oil, coal, electricity, labour and freight[55]) and flat in 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, Z_{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 supply-shifter may be viewed as an additional element of Z_{lt}.) The market-level pricing equation is thus

$$p_{lt} = -\theta_l \frac{p_{lt}}{\hat{\eta}_{lt}} + W_{lt}\beta^1 + Z_{lt}\beta^2 + \nu_l + \varepsilon^s_{lt} \tag{14}$$

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 stabilisation). Equation (14) 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 an estimate based on the demand estimates from Section 4.2, 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}, Z_{lt}) and the estimates of the fixed coefficients (β, ν) as $W_{lt}\hat{\beta}^1 + Z_{lt}\hat{\beta}^2 + \hat{\nu}_l$.

Figure 15 reports estimation results[56]. The coefficients on the prices of fuel oil, coal, electricity and freight are all positive and significant. The coefficient on the average size (age) of plants is negative (positive), as expected, though not significant. On the other hand, contrary to intuition, the price of labour is significantly negative[57]. The price-cost margins are estimated to be very low; these are pictured in Figure 16, 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[58]. 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 14, the estimated (average) price-cost margin as a proportion of price is only $0.0079/0.48 \approx 1.6\%$ (recall expression (2)).

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[55] 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 modelled as the 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).

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

[57] One could, however, attempt to rationalise this result by reference to a rent-sharing story. See, for example, Clark (1980).

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

Figures 15 and 16 about here

It is clear from our knowledge of marginal cost and price-cost margins in the industry that these estimates are inconsistent. Figure 16 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 (Proposition 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.[59]

4.4 Inferring conduct in a constrained equilibrium

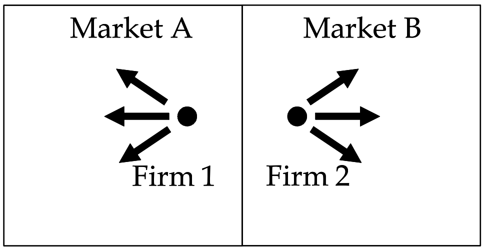
Prior to stating the results of the test of Proposition 3 as it applies to the Brazilian cement industry, I provide a flavour of why the data leads to the rejection of Cournot behaviour by considering a specific example extracted from the data. This serves only as an example of a broader trend. As shown subsequently, there are many instances in the data where firms undersupply local markets as compared to the supply decisions of a Cournot firm.

A case in point: 2 firms, 2 markets, 1 time period Consider the two adjacent states of Alagoas (*AL*) and Sergipe (*SE*), located in the northeast of Brazil (see Figure 6). These states are equally small both in terms of market size and geography. Up until 1996 each was home to only one plant: the firm Brennand 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)[60]. Consider the year 1996. While firm 1 commands an 83% share in market A, it does not supply to neighbouring market B,

[59] A comment on this particular industry where conditions such as the steepness of the demand curve result in demand being inelastic at the equilibrium. Assume one overlooks the binding imports constraint and thus misguidedly considers the class of behavioural 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 behaviour ($\theta = 1$): 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 a Cournot industry, unless all firms have small market shares. To see this, notice that if the largest firm has a market share of 50%, then $\max_f\{\theta_f\} = 0.5$ and thus this firm's marginal revenue, to be equated to its marginal cost, is zero. As seen, 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 model and the Cournot model being rejected in favour of price-taking behaviour 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%!

[60] In late 1996 a third firm, GJS, set up a plant close to Votorantim's plant in *SE*. However, I abstract away from this in this illustration, by considering the year 1996.

right next door to its plant located in market A. Equally striking, firm 2 commands an 89% share in market B, while attaining only a 7% share in the neighbouring market A, next door to its plant in market B. Average consumer prices in markets A and B are almost identical, respectively R$ 9.47 (per bag) and R$ 9.46. As explained previously, 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$ 5.17.[61] This is illustrated in the following picture and table, where I take the price elasticity of demand in equilibrium for both markets to be -0.5:[62]



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

(Year 1996)	Price, p	Share, $\frac{q_f}{q}$	MR Cournot, $p+\frac{p}{\eta}\frac{q_f}{q}$	MC, c	Can reject Cournot?
Local market $l=A$ (AL)					
Firm $f=1$ (Bren)	9.47	0.83	$9.47-\frac{9.47}{0.5}0.83=-6.25$	5.20	No ($\varphi_{1A}<0$)
Firm $f=2$ (Voto)	9.47	0.07	$9.47-\frac{9.47}{0.5}0.07=8.14$	5.30	**Yes** ($\varphi_{2A}>0$)
Local market $l=B$ (SE)					
Firm $f=1$ (Bren)	9.46	0	$9.46-\frac{9.46}{0.5}0=9.46$	5.47	**Yes** ($\varphi_{1B}>0$)
Firm $f=2$ (Voto)	9.46	0.89	$9.46-\frac{9.46}{0.5}0.89=-7.38$	5.17	No ($\varphi_{2B}<0$)

One is thus able to reject the hypothesis of firm 1 behaving in Cournot fashion towards market B in 1996, since (perceived) marginal revenue $p+\frac{p}{\eta}.\frac{q_f}{q}=9.46$ considerably exceeds marginal cost 5.47, i.e. $\varphi_{1B}=9.46-5.47=3.99$, amounting to 42% of consumer price. Likewise, I reject Cournot behaviour for firm 2 towards market A in 1996: marginal revenue 8.14 considerably exceeds marginal cost 5.30, i.e. $\varphi_{2A}=2.84$, or 30% of consumer price. Firm 1's (firm 2's) supply decision toward market B (market A) corresponds to that of firm f in the right panel of Figure 5 in the theoretical framework of Section 2. A story where Votorantim tacitly agrees to give Brennand the upper hand in *AL* in exchange for the latter staying away from *SE* would help explain the observed shipments[63].

[61] 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$ 0.13. 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 penalised 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*.

[62] The subsequent test takes the estimation of η in Section 4.2 into account. Also for simplicity, in this illustration I compute an average φ for the year (for each firm-market combination).

[63] As noted previously, with a view to testing conduct, the marginal costs I construct are conservative, i.e. they err on the high side. (This understates φ, working against the rejection of Cournot conduct.) In spite of this, the φ are not only positive but sizeable: of the order of 30 - 40% of consumer price! Interestingly, note that Brennand (firm 1) ships from its plant in *AL* (market A) to the states of *PB*, *PE* and *BA*, located at further distances than *SE* (market B) and where prices are similar to those in *SE*.

I now compute the test statistic of Proposition 3 for each active firm-market-month combination, (f, l, t). A firm is active in a given month if it owns a plant which is active in that month; that is, 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, 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. Now at each month t 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. $c_{flt} := \min_{i \in \mathcal{O}_{ft}} c_{ilt}$. From the Cournot pricing condition (10), compute the test statistic

$$\hat{\varphi}_{flt} = p_{lt} + \frac{p_{lt}}{\hat{\eta}_{lt}} \frac{q_{flt}}{q_{lt}} - c_{flt} \tag{15}$$

where $\hat{\eta}_{lt}$ is based on the demand estimates of Section 4.2 and p, q and c are observed, such that the randomness in $\hat{\varphi}$ stems from the randomness of the estimated price elasticity $\hat{\eta}_{lt}$. Recall that $\hat{\varphi}_{flt} > 0$ is sufficient to reject the null hypothesis that firm f is behaving in Cournot fashion towards market l in month t, in favour of more collusive behaviour; importantly, this statistic allows for the constraining effect of imports.

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

The table above summarises the results. There are 37536 active firm-market-month combinations (corresponding, therefore, to an average of $37536/17/156 \approx 14$ active firms across the country in any given month). Since I calculate the 95% confidence interval (C.I.) for φ_{flt} from the 95% C.I. for the price elasticity η_{lt}, I choose to drop 12840 observations for which the upper limit to the C.I. for η_{lt} is positive. In other words, I conservatively consider only the 24696 combinations for which the C.I. for the price elasticity lies in the interval $(-\infty, 0)$. I find that the null hypothesis of Cournot behaviour allowing for the constraining effect of imports, $\varphi_{flt} \leq 0$, can be *rejected at the 5% level of significance in 14849 of these 24696 combinations*. In other words, under the Cournot conjecture, one would expect firms to expand their supply to local markets in $14849/24696 \simeq 60\%$ of monthly supply decisions vis-à-vis observed outputs – 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![64]

[64]To ensure robustness, Salvo (2005) provides the number of supply decisions for which $\hat{\varphi}_{flt} > 0$ when $\hat{\varphi}$ is calculated using elasticities of -0.3, -0.5 or -0.7. The rejection of Cournot behaviour in favour of more collusive conduct is robust. It is clear from (15) that increasing the (absolute value of the) elasticity reinforces this result.

Further, the *Supplementary Section* shows how such a collusive arrangement may be sustained in a context where firms meet in different markets. In light of the different local market structures observed in the Brazilian cement industry, I provide examples of simple dynamic games which in equilibrium give rise to the pattern of market division just identified.

5 Concluding remarks

This paper shows that the standard methodology for inferring supply underestimates 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 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 signalling, 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 behaviour 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: Shifts and rotations in the linear demand example

Begin by considering a shift in the demand curve as depicted in the left panel of Figure 3. Say the inverse linear demand curve $p = a - bq$ shifts outward to $p = a' - bq$, where $a' - a = da > 0$. Recall that under both alternative hypotheses of conduct – low-cost cartel constrained by imports, and high-cost competitive industry – the equilibrium shifts from E_1 to E_2, where $dp = 0$ and $dq = \frac{da}{b}$. Plugging this into the total derivative of the static pricing equation (5)[65] and noting that the demand slope $\frac{\partial p(q)}{\partial q}$ remains unchanged at $-b$, one obtains $0 = -\theta\left((-b)(\frac{da}{b}) + 0\right) + d\xi^s$. Thus

$$d\xi^s = -\theta da$$

from which it is clear that shifts in the demand curve are correlated with the error in the pricing equation (unless, of course, there is competition: $\theta = 0$). Now consider a rotation in the demand curve around E_1 (right panel of Figure 3). Say the inverse

[65] This may be written $dp = -\theta\left(\frac{\partial p(q)}{\partial q}dq + d(\frac{\partial p(q)}{\partial q})q\right) + d\xi^s$, considering θ and c are constant.

demand curve $p = a - bq$ rotates anticlockwise around $E_1 = (q_1, p_1)$ to $p = a' - b'q$, where $b' - b = db < 0$ and thus $a' - a = da = db.q_1 < 0$. Under both alternative hypotheses of conduct, the equilibrium remains stationed at E_1, and thus $dp = dq = 0$. Plugging this into the total derivative of the static pricing equation and noting that the change in the demand slope $d(\frac{\partial p(q)}{\partial q}) = -db$, one obtains $0 = -\theta\,(0 + (-db)q_1) + d\xi^s$. Recalling that $da = q_1 db$, this translates into

$$d\xi^s = -\theta q_1 db = -\theta da$$

so that rotations in the demand curve around the equilibrium are correlated with the error in the pricing equation.

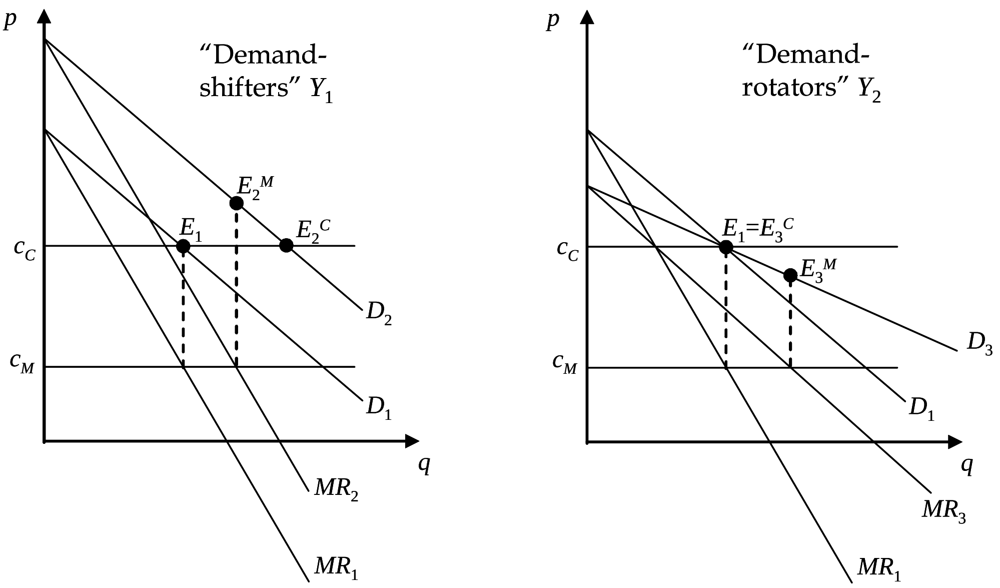

Figure 1: Identification in a static model. Left panel: Demand shifts. Right panel: Demand rotates.

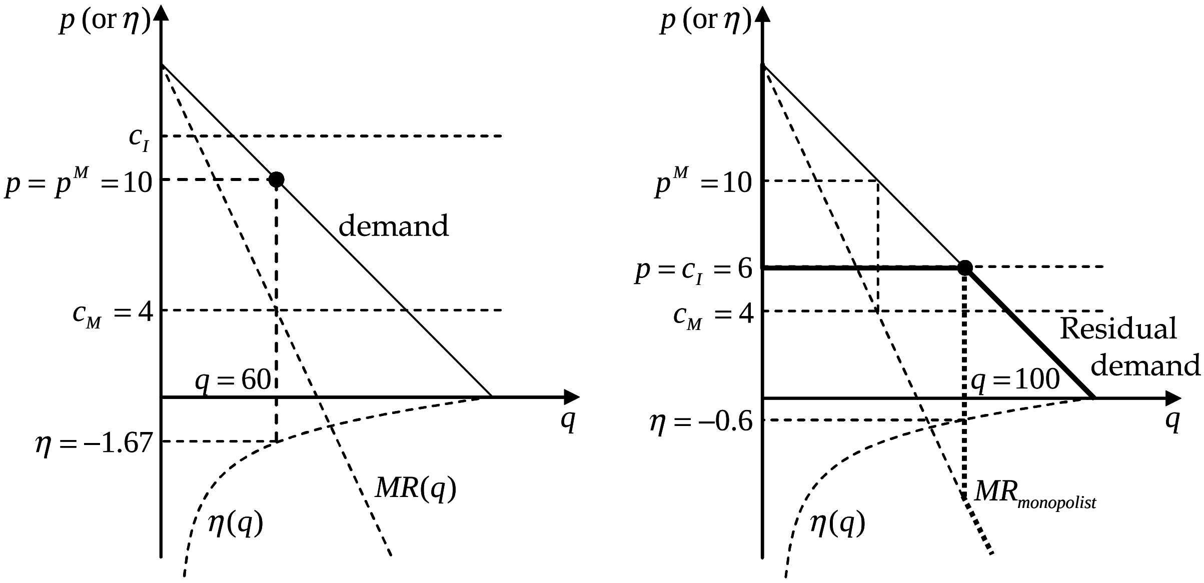

Figure 2: Monopolist facing a competitive fringe. Drawn for linear demand, given by $p = 16 - \frac{1}{10}q$ and $c_M = 4$. Left panel: Imports have no "bite" ($p^M < c_I$). Right panel: Imports constrain price in equilibrium ($p^M \geq c_I = 6$, as drawn).

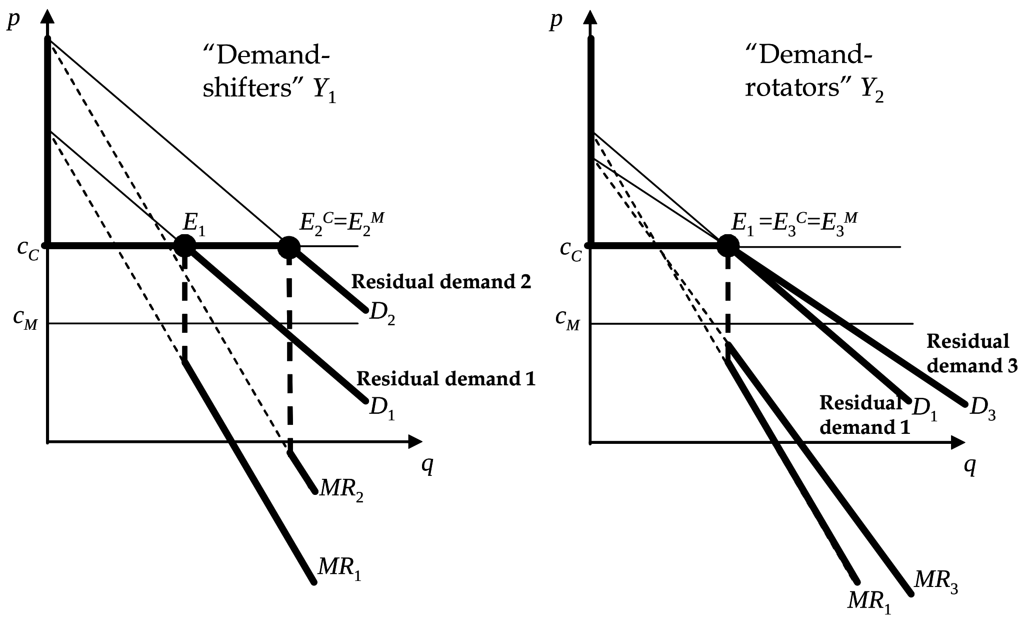

Figure 3: Conduct is no longer identified under the threat of entry. Left panel: Demand shifts. Right panel: Demand rotates.

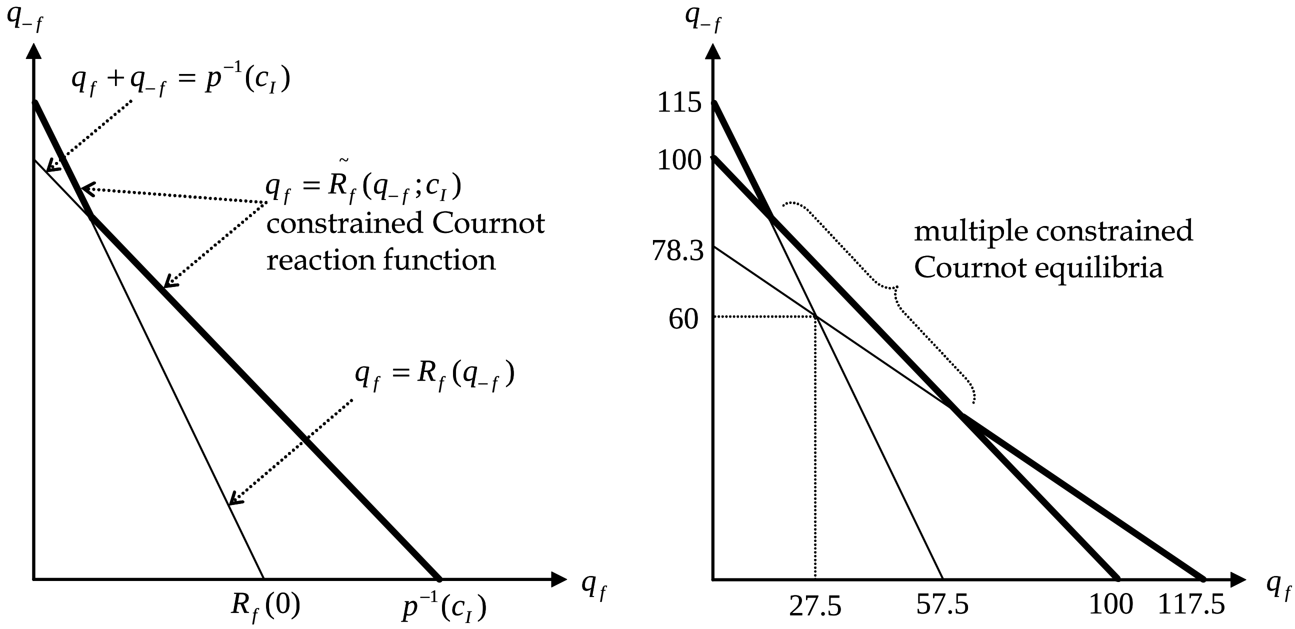

Figure 4: Cournot oligopoly facing a competitive fringe of imports. Left panel: Cournot firm f's reaction function, facing domestic rivals and imports. Right panel: Cournot equilibria. Drawn for linear demand $p = 16 - \frac{1}{10}q$ and $c_I = 6$, as in the constrained monopoly of Figure 2, now adding the assumptions that there are 3 firms, that the marginal cost of the firm of interest f is $c_f = 4.5$ and that $\sum_{j \neq f} c_j = 8.5$. Notice that joint output in the constrained Cournot oligopoly is 100, identical to that in the constrained monopoly. In the absence of imports, joint Cournot output would be 87.5.

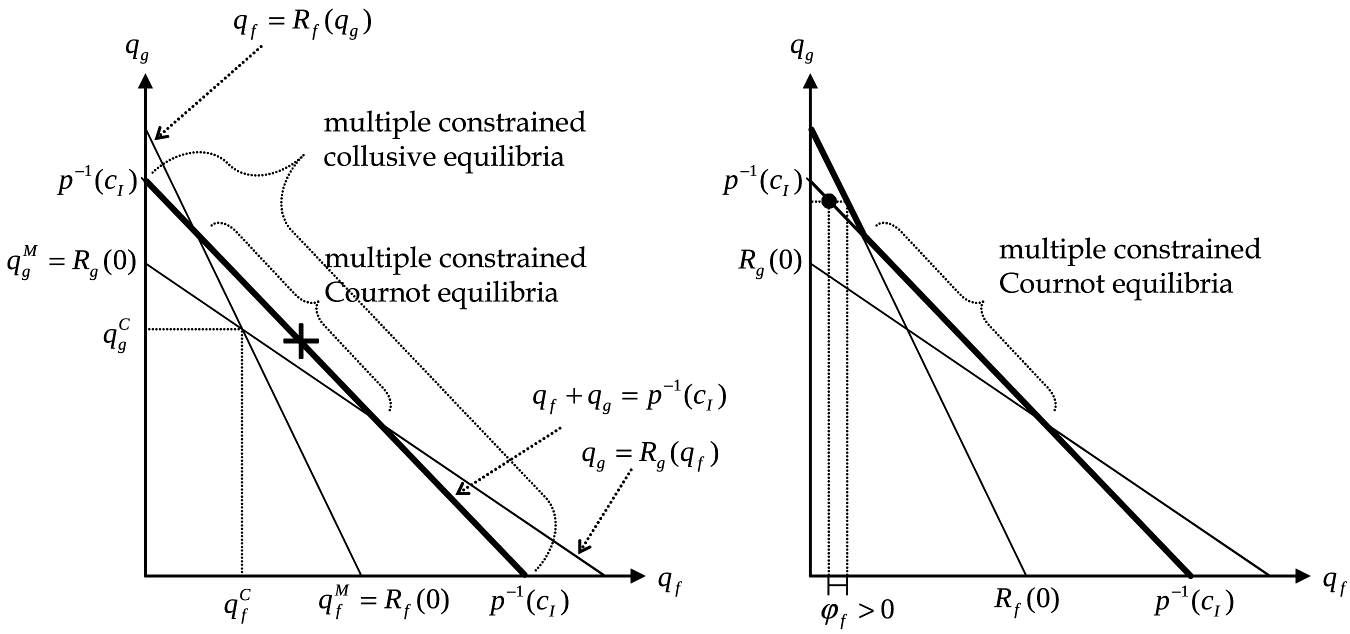

Figure 5: Identifying collusion from Cournot when imports constrain equilibrium prices under both models of conduct. Left panel: The imports constraint binds at the industry equilibrium marked "+", which is consistent with either Cournot conduct or more collusive firm conduct. Right panel: Rejection of Cournot behaviour for firm f: $\varphi_f > 0$.

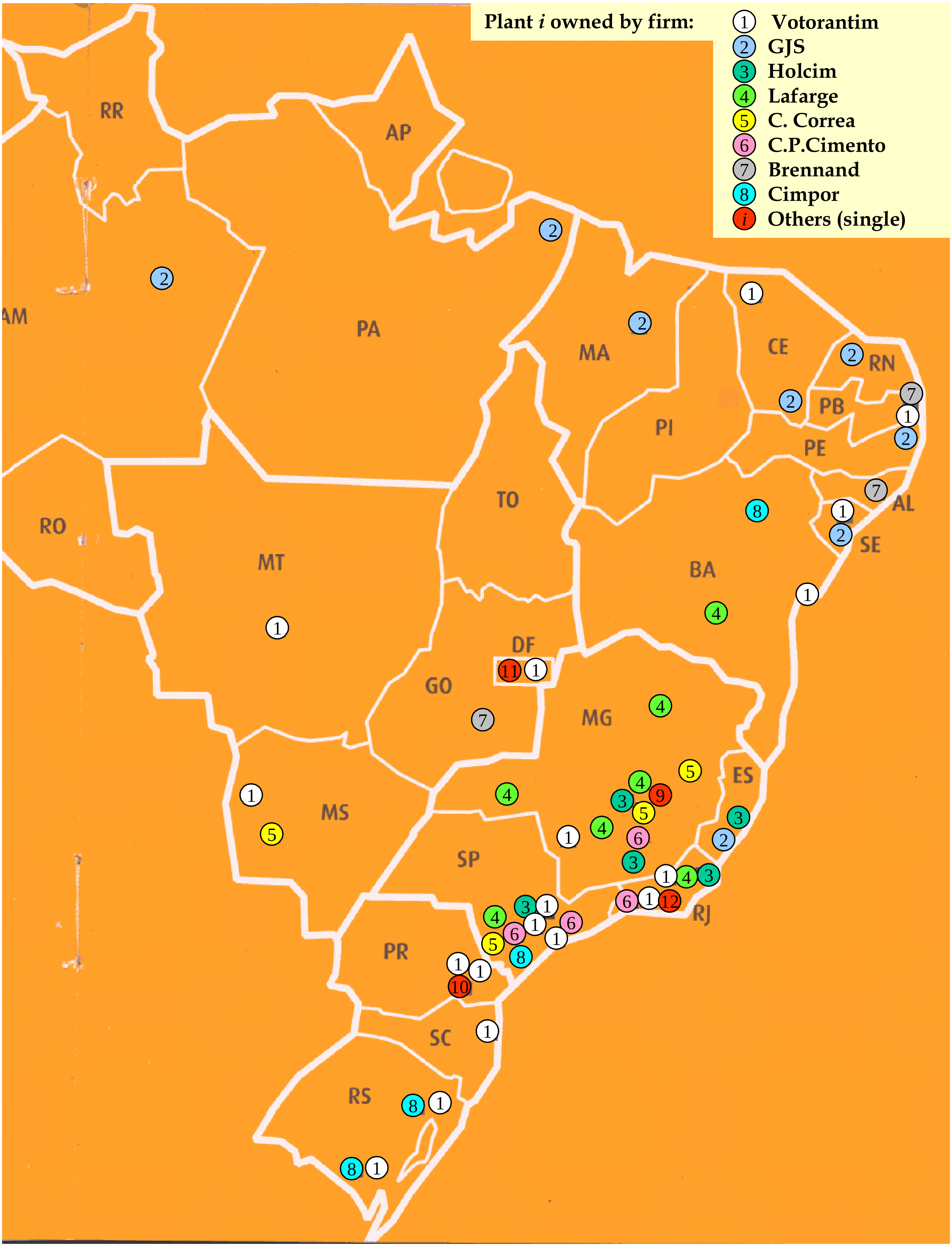

Figure 6: Active plants in 1999

	Mean	Standard Deviation	Maximum	Minimum	Memo: 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 kilometres)[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 7: Variation across 27 states of the Brazilian federation, Summary Statistics (time-varying figures refer to 1999)

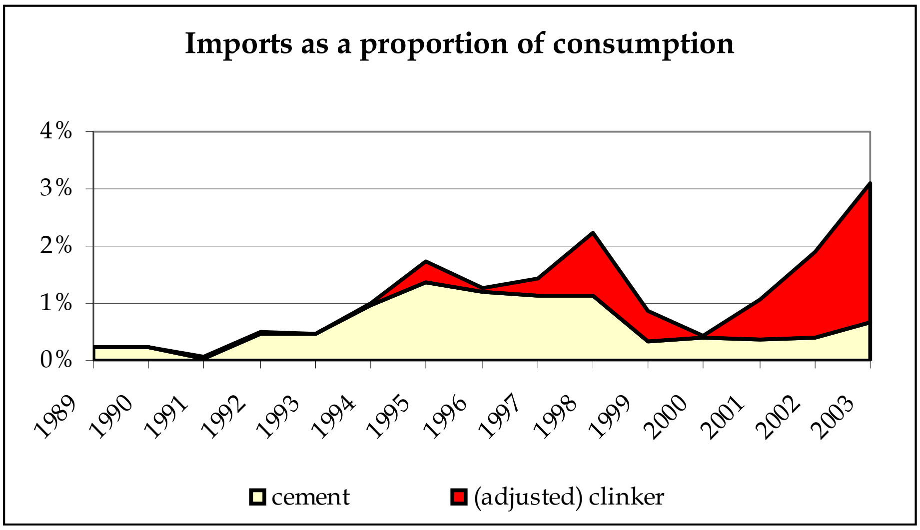

Figure 8: (Official) Imports of cement and clinker as a proportion of domestic consumption. Source: SECEX, MDIC. Clinker quantities are adjusted by the author to reflect usage in the production of cement (assumes 80% of clinker imports used in production of slag cement, with a 40% clinker content).

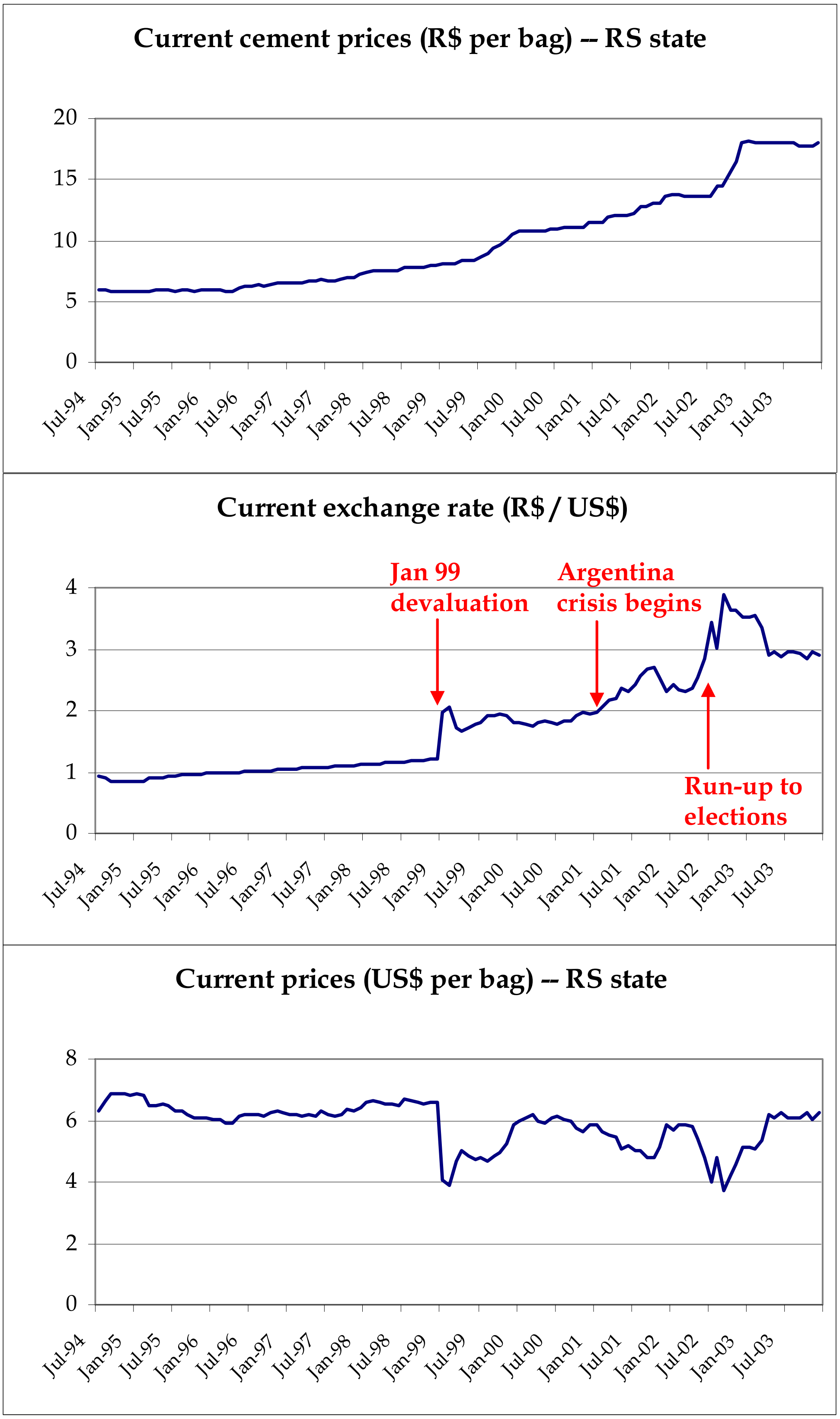

Figure 9: Evolution of cement prices in *RS* state since July 1994. In current local currency units (R$) per bag and US$ per bag

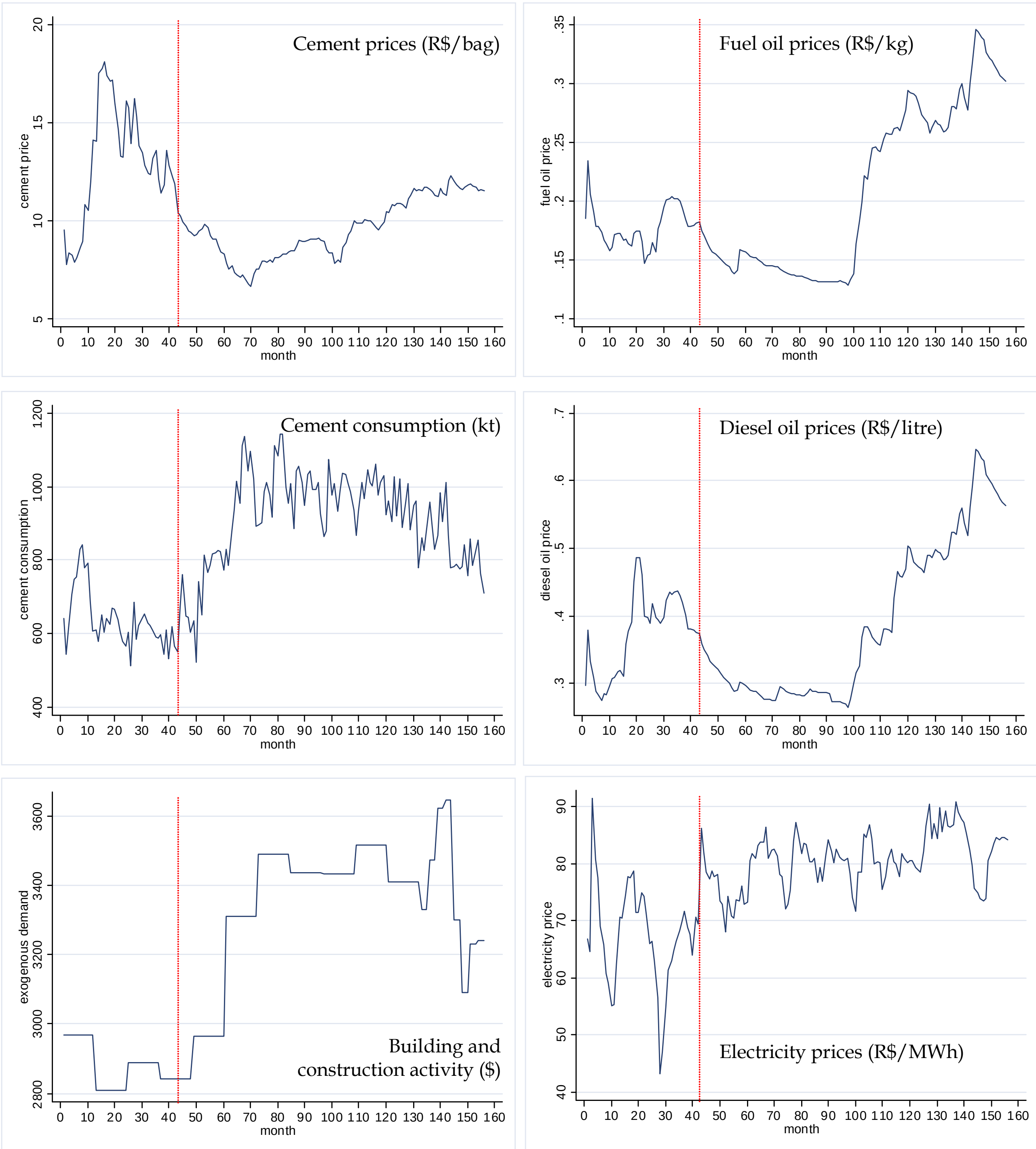

Figure 10: Cement prices, consumption, exogenous demand and factor prices for the state of São Paulo. All prices are in constant December 1999 values. Monthly observations, observation 1 corresponding to January 1991. July 1994, the month in which the stabilisation plan was enacted, is marked by the dotted lines.

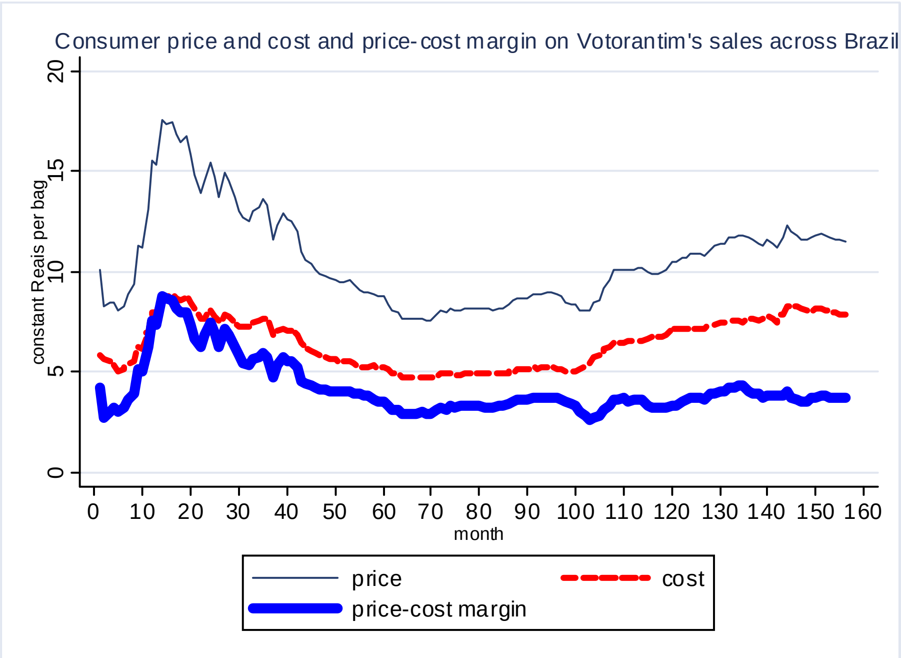

Figure 11: Evolution of consumer prices, marginal costs and price-cost margins on Votorantim's sales. Averaged across all states. In constant Reais per bag (at December 1999 values).

		(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-stabilisation:						
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-stabilisation:						
3,338	16% growth versus pre-stabilisation 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 12: Demand estimates for the state of *SP*

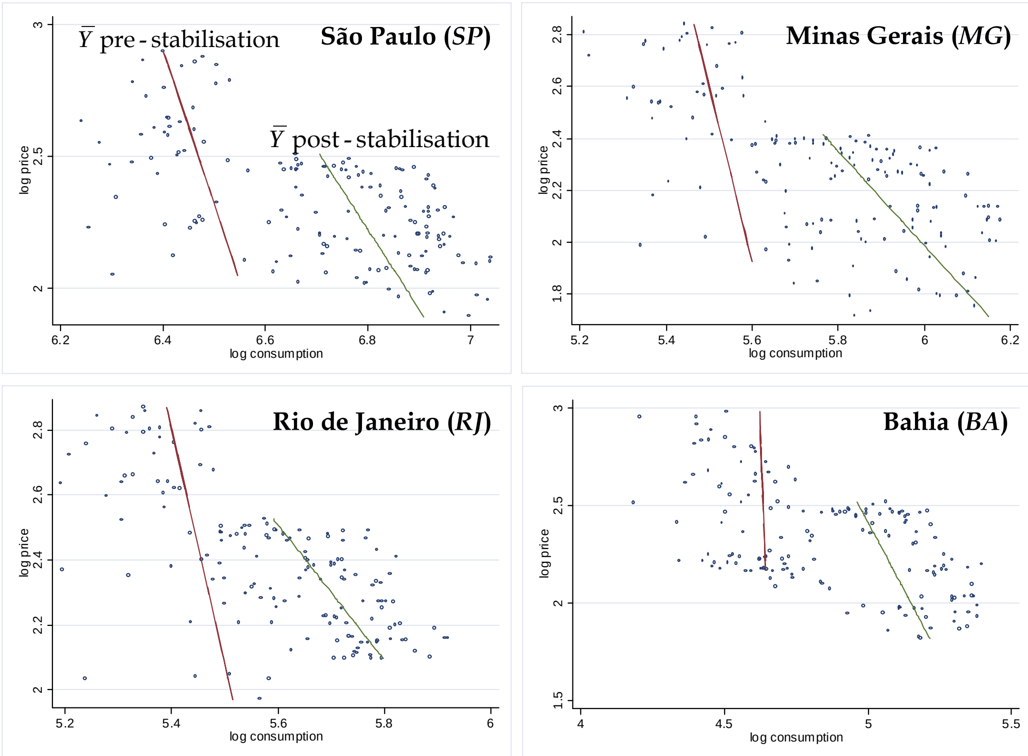

Figure 13: Fitted demand curves for the four largest markets. (Log) Price against (Log) Consumption. Evaluated at the respective means of exogenous demand Y for the pre- and post-stabilisation phases

State	Cement consumption in 1999 (kt)	(II) IV-imports bite: Interaction coef	s.e.	Log Price: Y evaluated at mean pre-stabilisation coef	s.e.	mean post-stabilisation 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 14: Demand estimates by state, Summary

	(I) IV Full sample		(II) IV Post-stabilisation 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)
Labour	-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-stabilisation 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 15: Estimation of a static pricing equation, assuming cost is not known. Instrumented with exogenous demand variables.

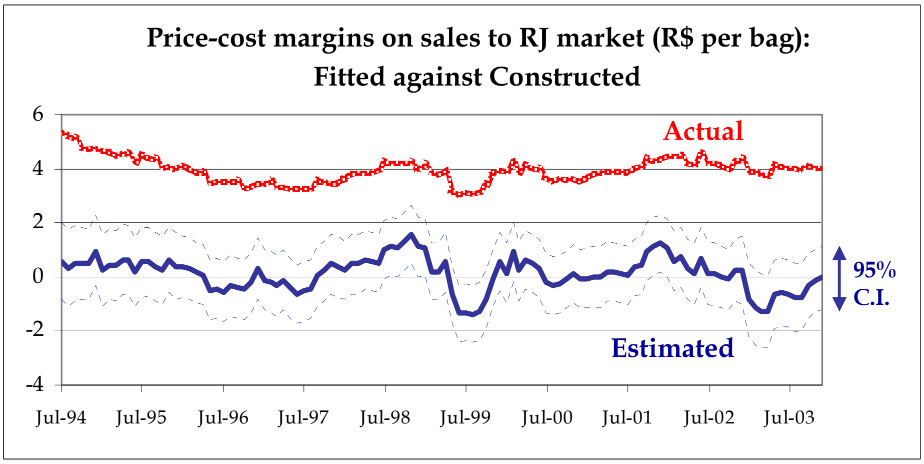

Figure 16: Estimated (average) price-cost margin on sales to RJ market, as estimated by the static pricing equation, against Actual (constructed) price-cost margin. In constant R$ per bag (December 1999 terms).