Capital-Market Imperfections and Countercyclical Markups: Theory and Evidence

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By JUDITH A. CHEVALIER AND DAVID S. SCHARFSTEIN*

During recessions, output prices seem to rise relative to wages and raw-material prices. One explanation is that imperfectly competitive firms compete less aggressively during recessions. That is, markups of price over marginal cost are countercyclical. We present a model of countercyclical markups based on capital-market imperfections. During recessions, liquidity-constrained firms boost short-run profits by raising prices to cut their investments in market share. We provide evidence from the supermarket industry in support of this theory. During regional and macroeconomic recessions, more financially constrained supermarket chains raise their prices relative to less financially constrained chains. (JEL E32, D43, G31)

Simple models of business cycles based on aggregate demand shocks imply that during booms, factor prices fall relative to output prices. This follows from the standard assumption that, at high output levels, marginal products are low. However, this implication is difficult to square with the facts. During booms, wages and raw-material prices tend to rise relative to output prices—that is, real fac-

tor prices are procyclical (see, for example, Gary Solon et al., 1994; and Kevin Murphy et al., 1989).

A number of papers have argued that imperfect competition can reconcile procyclical real factor prices with aggregate-demand-driven business cycles. These papers build on the old idea in Arthur C. Pigou (1927) and John Maynard Keynes (1939) that increases in aggregate demand may have little effect on prices—and thus a large effect on output—because firms behave more competitively during booms. As a result, output prices fall relative to marginal costs (that is, markups fall) and real factor prices rise.

There are at least three distinct reasons why markups may be countercyclical. First, demand may become less elastic during recessions, allowing imperfectly competitive firms to increase markups (see, for example, Mark Bils [1989], Paul Klemperer [1995], Arthur Okun [1981], and Joseph Stiglitz [1984] for reasons why elasticities of demand may be procyclical). Second, as argued by Julio Rotemberg and Garth Saloner (1986) and Rotemberg and Michael Woodford (1991, 1992), markups may be countercyclical because firms are less able to collude during booms. When demand is high, firms have greater incentive to cut prices because the short-run profits from stealing market share

are high relative to the long-run profits from collusion. Finally, Bruce Greenwald et al. (1984), Nils Gottfries (1991), and Klemperer (1995) have suggested that markups may be countercyclical because of capital-market imperfections. During a recession—when firms have low cash flow and greater difficulty raising external funds—they will try to boost current profits to meet their liabilities and finance investment. They may do so by increasing prices and forgoing attempts to build market share.

Our goal is to analyze the link between capital-market imperfections and countercyclical markups, and to test its empirical relevance. The starting point for our work is the large theoretical and empirical literature suggesting that information and incentive problems in the capital market can limit the ability of cash-constrained firms to make valuable investments.[1] We build on this literature by focusing on how liquidity constraints affect pricing behavior.[2] Just as capital-market imperfections can prevent firms from choosing investment projects that maximize the discounted value of profits, they can also prevent firms from choosing *prices* that maximize the discounted value of profits.

In Section I, we formalize the idea that liquidity constraints can affect pricing behavior. We start with a simple model of competition based on Klemperer's (1987) model of markets with consumer "switching costs" (see also Joseph Farrell and Carl Shapiro [1988]; Klemperer [1995] provides an overview of this approach). In this class of models, firms try to build market share by keeping prices down in the short run. Market share is valuable because consumers find it costly to switch

firms, and this gives firms market power over their repeat customers. This model can predict procyclical or countercyclical markups depending on the nature of the demand shocks and the parameters of the model.

However, capital-market imperfections and liquidity constraints tend to make markups countercyclical. We make this point with a model in which firms need to raise external funds to finance their operations. We model a particular type of incentive problem in which debt emerges as the optimal financial contract along the lines of Oliver Hart and John Moore (1989) and Bolton and Scharfstein (1996). Since firms may default, they have less incentive to build market share because they may not reap the benefits of the investment. During a recession, this effect is particularly strong because the probability of default is high. Thus, the model illustrates that capital-market imperfections combined with a market-share model of product-market competition can generate countercyclical markups.

We then empirically analyze the effects of capital-market imperfections on product-market competition in the supermarket industry. We study a single industry to avoid the problems associated with cross-industry comparisons of competition (see Timothy Bresnahan [1989] for a discussion of the limitations of such studies) a feature of existing studies of cyclical variation in markups. These difficulties also raise problems for the existing empirical studies of the movement of markups over the business cycle (see, for example, Rotemberg and Saloner, 1986; Rotemberg and Woodford, 1991, 1992; and Ian Domowitz et al., 1986). We study the supermarket industry in particular because firms compete in many local markets. This allows us to use price data for a cross section and time series while still examining a single industry.[3]

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[1] See, for example, Stewart Myers and Nicholas Majluf (1984) and Ben Bernanke and Mark Gertler (1989) for theoretical models along these lines, and Steven Fazzari et al. (1988) and Takeo Hoshi et al. (1991) for empirical evidence.

[2] Drew Fudenberg and Jean Tirole (1986) and Patrick Bolton and Scharfstein (1990) have also recognized that liquidity constraints can have product-market effects. Their point is that if firms cut investment when cash flow falls, rivals will have an incentive to ensure that cash flows are low. This theory can help to rationalize predatory practices because, at the extreme, cash constrained firms may completely disinvest (that is, exit the market).

[3] We are aware of one other study which estimates changes in markups in a single industry in local markets. Using the seasonal pattern of gasoline demand, Severin Borenstein and Andrea Shephard (1993) find that markups in retail gasoline markets are higher when demand in the near future is expected to be high. They interpret these results as consistent with models of tacit collusion such as Rotemberg and Saloner (1986). Since we would not

We present three pieces of evidence that suggest the importance of capital-market imperfections in generating countercyclical markups. The first is from examining local price changes during the severe recession that occurred in oil producing states as a result of the halving of oil prices in 1986. In some cities, national supermarket chains have large market shares, while in others, local and regional chains have a larger presence. We would expect the liquidity of the local and regional chains to be more adversely affected by the downturn since national chains also have operations in non-oil states that were performing well during this period. Thus, if there are capital-market imperfections and firms price for market share, we would expect prices to fall less (or rise more) in cities where local and regional chains have a large presence. Indeed, they do.

The second piece of evidence comes from examining the local price responses to the macroeconomic recession of 1990–1991. During the latter half of the 1980's many supermarket chains undertook leveraged buyouts (LBOs) which increased their debt ratios dramatically. We would expect these firms to be more liquidity constrained in response to an adverse economy-wide shock. Therefore, they should cut prices less (or raise them more) during this downturn. This seems to be the case, particularly in cities that did very poorly in the recession.

The third piece of evidence is from an examination of firm-level pricing in the period following the macroeconomic recession of 1990–1991. According to the theory, LBO firms should not cut prices as much as their less leveraged rivals in cities which continue to perform poorly following the macro recession. We would also expect firms to cut prices less in poorly performing cities if their rivals are highly leveraged. We find evidence of both of these effects.

The empirical results are consistent with our model of product-market competition in which firms price for market share and in which li-

quidity constraints affect pricing behavior. The results are inconsistent with the tacit collusion models proposed by Rotemberg and Saloner (1986) and Rotemberg and Woodford (1991, 1992). As discussed above, these models predict that in booms there is a greater temptation to deviate from the collusive outcome by cutting prices in an attempt to increase short-run profits. However, adding liquidity constraints to their model tends to reverse the prediction of countercyclical markups. If firms are more liquidity constrained in recessions, then they will be *more* tempted to cheat on a collusive arrangement because they need to increase short-run profits. Thus, their model predicts that prices should fall *more* in busts when firms are most cash constrained. By contrast, we find that prices fall *less*.

Although our principal focus is on countercyclical markups, the paper adds support to recent work on the link between capital markets and product markets. Chevalier (1995b) also looks at the effects of liquidity on supermarket pricing. The paper shows that following an LBO—an event which reduces corporate liquidity—local supermarket prices tend to rise if there are already many other LBO firms in the market. We are taking the same basic approach, only studying different events that reduce liquidity.[4]

The paper is organized as follows. In Section I we outline a simple model which shows how capital-market imperfections can generate countercyclical markups. In Section II we begin discussion of our empirical approach. Section III presents the results from the oil shock and Sections IV and V analyze the effect of leverage on pricing during and following the 1990–1991 recession. Section VI concludes the paper with a discussion of the results and related research.

I. The Model

In this section, we present a simple model of the effect of capital-market imperfections

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expect liquidity to vary over the predictable seasonal demand cycle, this evidence does not bear on the question of how liquidity affects markups.

[4] Chevalier (1995a) also finds that non-LBO firms are more likely to enter and expand in local markets where LBO firms have a significant market share. This is consistent with the view that these are more attractive markets for investment because LBO firms raise prices.

on price-cost markups. As a benchmark, we start with a model in which firms are financed with internally generated funds. We then introduce external financing in an imperfect capital market and compare the equilibria.

The basic model of product-market competition builds on Klemperer (1996). Two firms, A and B compete for two periods, $\tau = 1, 2$. Their marginal cost of production is constant and equal to c_τ in period τ. Consumers have a reservation value of R for each unit they purchase. They are distributed with uniform density on the line segment, [0, 1], with firm A located at 0 and firm B located at 1. In the first period, they bear a "transportation cost" of t per unit of distance traveled along the line to the firm of their choice. Thus, a consumer located at $y \in [0, 1]$ would incur a cost of ty to buy from A and $t(1 - y)$ to buy from B. One can take this transportation cost at face value, as might be reasonable in retail markets (such as supermarkets). Alternatively, one can view this cost as stemming from nonspatial product differentiation: it measures how far each firm's product is from a consumer's ideal set of product characteristics. For simplicity, we assume that transportation costs are zero in the second period. However, we make the key assumption that in the second period consumers incur a switching cost, s, to buy from a different firm.

Because we are ultimately interested in how changes in demand affect equilibrium prices, we allow demand to vary across the two periods. Expected demand in the first period is $\bar{\theta}$ and expected demand in the second period is normalized to 1. For each firm, first-period demand can be high (θ_H) with probability μ or low (θ_L) with probability $1 - \mu$. We interpret a high value of μ as a boom and a low value of μ as a bust. While the value of μ is the same for both firms, the actual realization of demand, θ_H or θ_L, is firm specific. We think of this firm-specific demand realization as stemming from some unmodeled aspect of a firm's product that affects demand.[5] Firms choose prices before they know this demand realization.

Finally, in order to compete in this market, firms must invest an amount I at the beginning of the first period.

We first solve for the equilibrium in the second period. As noted above, after purchasing from a firm in the first period, the consumer incurs a switching cost, s, to buy from the other firm. If this switching cost is high enough, each firm can charge the consumer's reservation price, R, without fear of being undercut by its rival because the rival would have to cut the price a discrete amount to $R - s - \varepsilon$. While the rival may sell more units at this lower price, it earns considerably less on its locked-in first-period customers.

It follows that second-period profits for each firm, $k = A, B$, depend on their first-period markets shares, σ_1^k. In particular, we can write firm k's second-period profits as

$$(1) \qquad \pi_2^k(\sigma_1^k) = (R - c_2)\sigma_1^k.$$

In the first period, firms take into account that their second-period profits are higher if they capture more of the market in the first period. If firm A charges p^A and firm B charges p^B in the first period, a consumer located at point y will buy from A rather than B provided

$$(2) \qquad p^A + ty \le p^B + t(1 - y),$$

or

$$(3) \qquad y \le \frac{1}{2} + \frac{p^B - p^A}{2t}.$$

Consumers will buy from B if the inequality is reversed. Note that inequality (3) is only valid if price plus transportation cost is less than the consumer's reservation value, R. We assume that the parameters are such that this condition is met.

From (3) it follows that the market shares of firm A and B in period 1 are given by

$$(4) \qquad \sigma_1^A = \frac{1}{2} + \frac{p^B - p^A}{2t} = 1 - \sigma_1^B.$$

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[5] The assumption that demands θ_H and θ_L are firm specific will later eliminate any need to consider financial contracts which are dependent on the other firm's actions.

Thus, first-period profits for firm A in state i can be written as

$$(5) \quad \pi_1^A(p^A, p^B, \theta_i) = (p^A - c_1)\theta_i\sigma_1^A$$

$$(6) \quad = (p^A - c_1)\theta_i\left[\frac{1}{2} + \frac{p^B - p^A}{2t}\right],$$

and analogously for B.

Firm A chooses p^A given its conjecture about p^B to maximize the discounted value of its profits over two periods. Assuming the discount rate is 0, this amounts to maximizing

$$(7) \quad (p^A - c_1)\bar{\theta}\sigma_1^A + (R - c_2)\sigma_1^A.$$

The first-order conditions for this problem define firm A's pricing reaction curve as a function of firm B's price:

$$(8) \quad p^A = \frac{t + c_1}{2} + \frac{p_1^B}{2} - \frac{R - c_2}{2\bar{\theta}}.$$

As is standard, firm A's optimal price is increasing in its rival's price; in the terminology of Jeremy I. Bulow et al. (1985), prices are "strategic complements."

We write the equilibrium price for firm A as $p^{*A}(e_A, e_B)$ and for firm B as $p^{*B}(e_A, e_B)$, where $e_k = 1$ if firm k needs external financing and $e_k = 0$ if firm k does not need external financing. Thus, the symmetric equilibrium price, $p^*(0, 0)$, when both firms are internally financed is

$$(9) \quad p^*(0, 0) = t + c_1 - \frac{R - c_2}{\bar{\theta}}.$$

The markup of price over marginal cost for firm i, which we write in analogous fashion as $m^{*i}(e_A, e_B) = p^{*i}(e_A, e_B) - c_1$, in this case is just $m^*(0, 0) = p^*(0, 0) - c_1 = t - (R - c_2)/\bar{\theta}$.

In a one-period model, each firm would charge a price of $t + c_1$. However, here prices are less than $t + c_1$ because firms compete for first-period market share so that they can later charge the monopoly price, R, to their locked-in customers. This effect lowers prices by $(R - c_2)/\bar{\theta}$. In fact, this effect can be so strong

that firms charge prices below marginal cost in the first period.

In this model, markups are procyclical. If we think of μ as measuring the level of demand, then given that price is additive in marginal cost, markups will be procyclical if $dp^*(0, 0)/d\mu > 0$, and countercyclical if the inequality is reversed. Differentiating (10) with respect to μ we have

$$(10) \quad \frac{dm^*(0, 0)}{d\mu} = \frac{dp^*(0, 0)}{d\mu}$$

$$= [R - c_2]\frac{\theta_H - \theta_L}{\bar{\theta}^2} > 0.$$

Markups tend to rise during booms because the increase in current demand relative to future demand makes it less attractive to price low to increase future monopoly profits. It is better to reap profits in the first period by charging a high price when demand is relatively high.[6]

Capital-market imperfections tend to reverse this basic result, making markups countercyclical or at least less procyclical. There are many ways to introduce capital-market imperfections. We choose a variant of Bolton and Scharfstein (1990, 1996) and Hart and Moore (1989) because it allows us to analyze optimal financial contracts in a simple model. The basis of these models is the assumption that corporate cash flow is observable to the manager and investors, but is not "verifiable"—that is, it cannot be observed by outside parties and therefore contracts cannot be made contingent on its realization. Moreover, the manager can, if he chooses, costlessly divert all of the cash flow to himself. This formulation captures the notion that managers can spend corporate resources on perks, pet projects, and so on, and that such spending cannot be directly controlled through contractual means.

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[6] Klemperer (1995) has shown that in a more general version of the model markups could be countercyclical. If, during booms, more new customers enter the market, then there will be more incentive to price low to attract these customers. Provided this effect is large enough, markups could be countercyclical. However, liquidity constraints will make markups even more countercyclical.

The only way to get managers to pay out cash flow is to threaten to liquidate the firm's assets if they do not. However, liquidation is inefficient in that the firm's assets are worth less if owned and managed by the investors. In particular, the firm's assets are worth a fraction, $\lambda < 1$, of the remaining cash flow. Thus, for example, if firm k is liquidated at date 1, it would be worth $\lambda \pi_2^k(\sigma_1^k)$ in the hands of the investor.

As Hart and Moore (1989) and Bolton and Scharfstein (1996) have shown, the optimal contract calls for a repayment of D at date 1; if no such payment is made, the investor has the right to seize the project's assets. This contract gives the manager some incentive to pay out D if he has the cash to do so. If the manager fails to pay, the asset may be liquidated and he loses the ability to divert cash to himself at date 2.

If the assets have not been seized at date 1, then the investor has no further leverage over the manager and he will therefore divert all of the period-2 cash flow of $\pi_2^k(\sigma_1^k)$ to himself. Working backwards, this means that the maximum payment than can be extracted from the manager at date 1 is $\pi_2^k(\sigma_1^k)$; incentive compatibility requires that $D \leq \pi_2^k(\sigma_1^k)$.

If, however, the manager does not have enough cash to make the payment D, say because $\pi_1^k(\sigma_1^k) < D$, then the manager would instead choose to pay nothing and have the asset liquidated. His first-period payoff would therefore be $\pi_1^k(\sigma_1^k) < D$ and he would receive no cash in the second period.

To make the problem interesting, suppose that in high-demand states cash flow is enough to cover the payment D and in low-demand states cash flow is less than D, that is, $\pi_1^k(\theta_H) > D > \pi_1^k(\theta_L)$. The manager's expected payoffs over the two periods, V^k, can be written as

(11) $V^k = \mu[\pi_1^k(\theta_H) - D + \pi_2^k(\sigma_1^k)]$

$+ (1 - \mu)\pi_1^k(\theta_L)$.

The investor will be willing to lend provided his expected payoffs are nonnegative:

(12) $\mu D + (1 - \mu)\lambda \pi_2^k(\sigma_1^k) - I \geq 0$.

Competition among investors ensures that condition (12) is met with equality. Note that D is chosen taking into account the product-market equilibrium that follows in periods 1 and 2. However, if this value of D, $D^* = [I - (1 - \mu)\lambda \pi_2^k(\sigma_1^k)]/\mu$, is greater than $\pi_2^k(\sigma_1^k)$, then the contract would not be incentive compatible and there would be no feasible contract. We assume for the remainder that the parameters are such that $D^* \leq \pi_2^k(\sigma_1^k)$, that is, incentive compatible contracts are feasible.

Suppose that firm A needs to raise I externally. After the financial contract is signed, firm A chooses p^A taking D and p^B as given. The first-order condition for p^A is

(13) $\dfrac{\partial V^A}{\partial p^A} = \mu \left[\dfrac{\partial \pi_1^A(\theta_H)}{\partial p^A} + \dfrac{\partial \pi_2^A(\sigma_1^A)}{\partial p^A} \right]$

$+ (1 - \mu) \dfrac{\partial \pi_1^A(\theta_L)}{\partial p^A}$

(14) $= \bar{\theta} \left[\dfrac{1}{2} + \dfrac{p^B}{2t} + \dfrac{c_1}{2t} - \dfrac{p^A}{t} \right]$

$- \mu \dfrac{R - c_2}{2t} = 0.$

The first-order condition (14) resembles the first-order condition with internal financing with one exception. With internal financing, firms are never liquidated and they receive all of the second-period profits. Therefore, a marginal increase in p^A reduces second-period profits by $(R - c_2)/2t$. However, with external financing, the firm is liquidated with probability $1 - \mu$, so that the manager only gets second-period profits with probability μ. Thus, a marginal increase in p^A only reduces second-period profits by $\mu(R - c_2)/2t$. This means that for each p^B, firm A will charge a higher price when it is externally financed than when it is internally financed. With external financing, firms are less interested in building market share because they get less of the gains from doing so; in effect, they are more "short-term oriented."

Thus, firm A's reaction curve is given by

(15) $p^A = \dfrac{t + c_1}{2} + \dfrac{p^B}{2} - \mu \dfrac{R - c_2}{2\bar{\theta}}.$
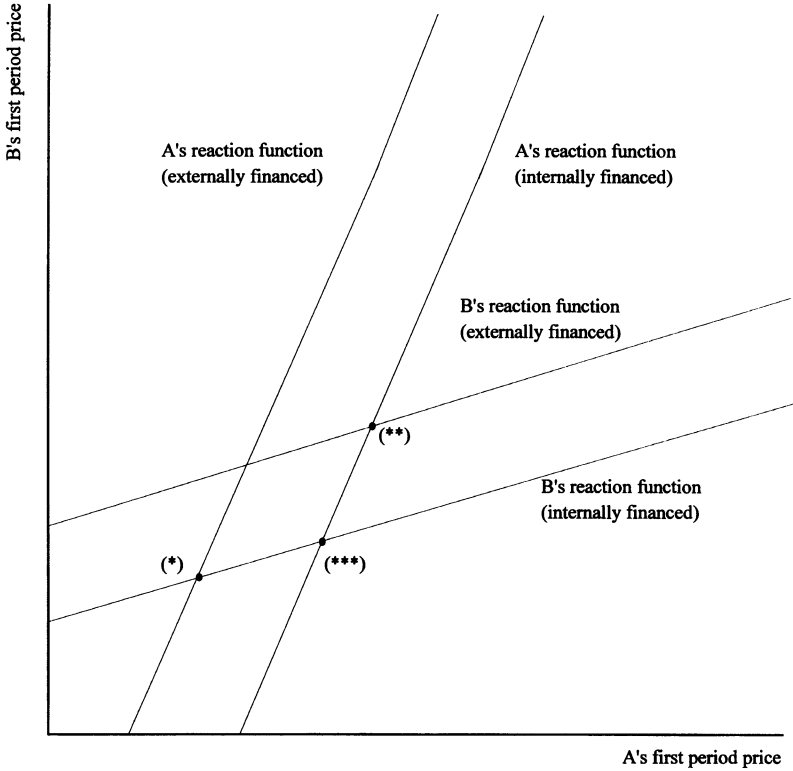
FIGURE 1. REACTION CURVES FOR FIRMS A AND B
WITH INTERNAL AND EXTERNAL FINANCING

If both firms need external financing, then the symmetric equilibrium price, $p^*(1, 1)$ is given by

$$(16) \quad p^*(1, 1) = t + c_1 - \frac{\mu}{\theta} (R - c_2).$$

Equations (9) and (16) imply that for all $\mu < 1$ the equilibrium price is higher when firms are externally financed than when they are internally financed.

One can see this in Figure 1 which shows both firms' reaction curves under the two financing regimes. Point (*) is the equilibrium when both firms are internally financed. If firm A is externally financed, its reaction curve shifts upward and to the left, while if firm B is externally financed its reaction curve shifts outward and to the right. This shifts the equilibrium to (**), at a higher price.

If only one of the firms, say firm A, is externally financed, then only its reaction curve shifts and the equilibrium is at (***). In this case, firm A's price is

$$(17) \quad p^{*A}(1, 0) = t + c_1 - \frac{R - c_2}{\theta}$$

$$+ \tfrac{2}{3}(1 - \mu) \frac{R - c_2}{\theta} \, .$$

Firm B's price is

$$(18) \quad p^{*B}(1, 0) = t + c_1 - \frac{R - c_2}{\theta}$$

$$+ \tfrac{1}{3}(1 - \mu) \frac{R - c_2}{\theta} \, .$$

In the general case, the equilibrium prices for firms A and B are

$$(19) \quad p^{*A}(e_A, e_B) = t + c_1 - \frac{R - c_2}{\bar{\theta}}$$

$$+ \frac{2e_A + e_B}{3}(1 - \mu)\frac{R - c_2}{\bar{\theta}}, \quad \text{and}$$

$$(20) \quad p^{*B}(e_A, e_B) = t + c_1 - \frac{R - c_2}{\bar{\theta}}$$

$$+ \frac{2e_B + e_A}{3}(1 - \mu)\frac{R - c_2}{\bar{\theta}}.$$

The first three terms of (19) and (20) sum to the price that A and B, respectively, would charge absent any financing constraints—it is the same expression as that given in (9). The last term in each equation reflects the effect of financing constraints on their prices.

A number of important conclusions emerges from equations (19) and (20). First, each firm's price is higher if it needs external financing than if it is internally financed: $p^{*A}(1, e_B) > p^{*A}(0, e_B)$ and $p^{*B}(e_A, 1) > p^{*A}(e_A, 0)$. Second, each firm's price is higher if the other firm is externally financed: $p^{*A}(e_A, 1) > p^{*A}(e_A, 0)$ and $p^{*B}(1, e_B) > p^{*A}(0, e_B)$.

Third, and most importantly, the degree of markup cyclicality depends on the financing behavior of both firms. For firm A,

$$(21) \quad \frac{dm^{*A}(e_A, e_B)}{d\mu} = \frac{dp^{*A}(e_A, e_B)}{d\mu}$$

$$= \left[\left[1 - \frac{2e_A + e_B}{3} \right] \theta_H - \theta_L \right]$$

$$\times \frac{R - c_2}{\bar{\theta}}.$$

The degree of markup cyclicality for firm B is given by an analogous expression.

Since the term in large brackets in (21) could be positive or negative, markups could be procyclical or countercyclical. However, the important point is that firm A's markup is more countercyclical (or less procyclical) if it is externally financed and if firm B is

externally financed. That is, (21) implies that

$$(22) \quad \frac{dm^{*A}(1, e_B)}{d\mu} < \frac{dm^{*A}(0, e_B)}{d\mu}$$

and

$$(23) \quad \frac{dm^{*A}(e_A, 1)}{d\mu} < \frac{dm^{*A}(e_A, 0)}{d\mu}.$$

In fact, if both firms are externally financed, the markup is countercyclical. In contrast, if both firms are internally financed, the markup is procyclical.

Note that since prices are strategic complements the financial position of a firm's competitor affects the cyclicality of the firm's own markup. Indeed, (21) implies that even if A is internally financed ($e_A = 0$) and thus faces no capital-market imperfections, its markup is countercyclical if B is externally financed ($e_B = 1$) and $\theta_H < {}^3/_2\theta_L$.

The model, narrowly interpreted, suggests that firms are less inclined to invest in market share during downturns because the increased probability of liquidation makes them care less about the future. However, the essence of the model does not depend on the extreme assumption of complete liquidation. In fact, it only requires that when firms have difficulty making debt payments they are unable to take full advantage of their locked-in customers. For example, they may be forced to scale back their operations or limit their expansion into related product lines.

We have shown that external financing leads to countercyclical markups because externally financed firms tend to increase markups when demand is low. The model takes as fixed whether or not a firm is externally financed. Yet, during a downturn, firms tend to rely more heavily on external financing as cash flow tends to fall faster than investment needs. Because externally financed firms have higher markups, an increase in the number of externally financed firms during a downturn will make markups even more countercyclical.

There are three empirically testable implications that emerge from the theoretical model.

1) A firm's markup should be more counter-cyclical if it is more financially constrained.

2) A firm's markup should be more counter-cyclical if its rivals are more financially constrained.

3) Average industry-wide markups should be more countercyclical if firms are more financially constrained.

We will first present evidence on industry-wide markups (Implication 3) in Sections III and IV and then present evidence on firm-level markups (Implications 1 and 2) in Section V.

II. The Empirical Approach

In our empirical analysis, we examine the effect of corporate liquidity on local-market pricing in the supermarket industry. The most direct approach to this issue would seem to be to relate firm-specific measures of the markup to measures of corporate liquidity. However, there are two difficulties with this approach, one conceptual and the other practical. The conceptual difficulty is that while liquidity may affect prices—the link explored in the model—prices almost certainly affect liquidity. This endogeneity problem makes it difficult to establish a causal link between prices and liquidity. The practical difficulty is that to measure markups one has to observe marginal costs. While a number of studies try to measure markups, they rely on strong assumptions about the form of the production function (see, for example, Bils, 1987; and Rotemberg and Woodford, 1992).

We take a different approach which we believe circumvents these endogeneity and measurement problems. We examine three exogenous events which reduced the liquidity of supermarket chains, but which were likely to have affected the liquidity of some chains more than others. We then investigate whether the chains for which liquidity was more adversely affected by these events raised their prices relative to the other chains.

First, we study supermarket prices during the recession in the oil-producing states brought on by the sharp drop in oil prices in early 1986. The liquidity of local and regional supermarket chains is likely to have been more ad-

versely affected by this shock than the overall liquidity of national supermarket chains since national chains have operations outside the oil-producing states. If liquidity affects pricing in the way predicted by our model, then supermarket prices should fall less in cities where local and regional chains have a large market share. (Implication 3 above).

We also examine supermarket pricing during the macroeconomic recession of 1990–1991. This is a useful episode to study because many supermarket chains undertook leveraged buyouts (LBOs) during the 1980's, which increased their leverage dramatically. We would expect the drop in liquidity brought on by the recession to have a larger impact on LBO firms than on firms that did not undertake LBOs. Thus, prices should fall less (or rise more) in local markets dominated by LBO chains (Implication 3 above).

Finally, we examine supermarket pricing in the period following the macroeconomic recession of 1990–1991, during which many states were still experiencing downturns. In this analysis we are able to compare the pricing behavior of LBO and non-LBO chains by using firm-specific pricing data at the local level. We examine whether LBO firms cut prices less than non-LBO firms (Implication 1) and whether firms cut prices less when their principal competitors are LBO firms (Implication 2).

There are two advantages of our approach. First, since the regional and macroeconomic recessions are exogenous, so are the shocks to liquidity. Therefore, one need not be concerned that the results are driven by an endogenous link between prices and liquidity.

Second, this approach does not require us to measure marginal costs, a feature of other attempts to measure markup variation over time. To see this more clearly, consider our study of the oil shock. By definition, the change in price from peak to trough, Δp, is just equal to the change in the markup, Δm plus the change in marginal cost, Δc. The difference between the price change for a local chain and the price change for a national chain is therefore

$$(24) \quad \Delta p^L - \Delta p^N = [\Delta m^L - \Delta m^N]$$

$$+ [\Delta c^L - \Delta c^N],$$

where the superscript L denotes local chains and the superscript N denotes national chains.

Equation (20) makes clear that prices of local chains could rise relative to prices of national chains for two reasons: (i) markups of local chains rise relative to markups of national chains; or (ii) marginal costs of local chains rise relative to marginal costs of national chains. Of course, effect (i) is the one in which we are interested. That is, if marginal costs of local chains rise relative to national chains, then we can draw no conclusion about the effect of liquidity on markups by looking at $\Delta p^L - \Delta p^N$.

While there is no particular reason to believe that the marginal costs of local chains rise relative to those of national chains, we can investigate this possibility by examining price changes outside the oil states. Our empirical strategy is, in effect, to compare $\Delta p^L - \Delta p^N$ in the oil states to $\Delta p^L - \Delta p^N$ in the non-oil states. Since the liquidity of neither the local nor the national chains were adversely affected in the non-oil states during this period, if we find that $\Delta p^L - \Delta p^N$ is essentially the same in the non-oil states as it is in the oil states, it would be difficult to argue that liquidity constraints increase the markups of local chains relative to the markups of national chains. We would have to conclude either that the marginal costs of local chains rose relative to the marginal costs of national chains, or that the markups of local chains rose relative to the markups of national chains for reasons quite apart from liquidity constraints.

However, if we find that $\Delta p^L - \Delta p^N$ in the oil states is greater than it is in the non-oil states, we can conclude one of two things: (i) markups of local chains rise relative to markups of national chains *only in the oil states*; or (ii) marginal costs of local chains rise relative to marginal costs of national chains *only in the oil states*. It is hard to think of any reason why interpretation (ii) would be true. But if local chains are more adversely affected by the oil-induced recession (and thus are more liquidity constrained), and if firms price for market share, then the markups of local chains should rise relative to the markups of national chains *only in the oil states*.

The same basic logic guides our investigation of the price behavior of LBO and non-LBO firms. First, we examine whether LBO firms raise prices relative to non-LBO firms in response to an economic downturn. If we find that they do, this could be because they raise markups relative to non-LBO firms, or because their marginal costs rise relative to non-LBO firms. Therefore, we also look at whether this effect is more pronounced in cities that were more adversely affected by the recession. If we find that LBO firms are more prone to raise prices relative to non-LBO firms in cities hit harder by the recession—that is, cities where liquidity constraints are more severe—then one could conclude that liquidity constraints increase the markups of LBO firms relative to non-LBO firms. The alternative interpretation is that the marginal costs of LBO firms rise relative to the marginal costs of non-LBO firms in cities that were more adversely affected by the recession, but it is difficult to see why this would be true.

While we think our approach circumvents endogeneity and measurement problems, it has some potential limitations. First, we are only examining a single industry so it is difficult to draw macroeconomic conclusions from our results. Second, this approach does not allow us to determine whether markups are counter-cyclical, only that liquidity constraints move markups in the direction of being counter-cyclical.

A third concern is that switching costs—the basis of the incentive to price for market share—may not be applicable to the supermarket industry. While there is no direct evidence of switching costs in this industry, there is some evidence that firms can raise prices in the short run without losing significant market share. D. Grant Devine and Bruce W. Marion (1979) report an experiment in which they placed advertisements in local newspapers in Ottawa-Hull, Canada, listing representative prices charged by local supermarkets. They find that many consumers switched supermarkets in response to the new information. This suggests that consumers have imperfect information about prices and continue to shop at a supermarket even if its prices are higher than competitors' prices.[7]

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[7] See Edmund Phelps and Sidney Winter (1970) for an

Thus it would seem that, as in our model with switching costs, firms can raise profits in the short run by increasing prices. This raises the question of why supermarkets themselves did not publish price comparisons. Unlike the researchers in this study, supermarket chains would have an incentive to publish information only on those products for which they had lower prices. Therefore, the published information would not credibly communicate overall price differences between supermarkets.

More evidence along this line is provided by Steven Hoch et al. (1993). They conducted an experiment with the cooperation of a Chicago-based supermarket chain in which they raised prices in some store locations for approximately 16 weeks. They found very low demand elasticities; a 10-percent price increase reduced sales volume by only 3 percent during this period. Since, on average, gross margins on supermarket products are 25 percent, one can show that a 10-percent price increase would increase gross profits by 36 percent in the short run. Given pretax net profit margins of about 2.5 percent, a 10-percent price increase would increase pretax net profits by more than four times.

So, this finding is also consistent with the basic feature of our model that an increase in the price increases short-run profits.

A final concern with our approach is that the industry may not be particularly cyclical.[8] While individual demand may not be very income elastic, demand at the local level may vary considerably with local economic conditions as individuals migrate into an area during booms and out of an area during busts. Indeed, this is probably why Olivier Blanchard and Lawrence Katz (1992) find that, at the state level, employment falls more than unemployment in recessions.

III. The Oil Shock

In this section we examine the regional downturn which occurred when oil prices dropped by about 50 percent in the first half of 1986. This caused severe recessions in states with relatively large production of oil and gas, particularly Texas, Louisiana, New Mexico, Oklahoma, Colorado, Wyoming, and Alaska. During this period, the rest of the economy was generally experiencing higher growth.

We examine the oil shock because it is the cleanest example of a regional shock. Other regional shocks include the farm crisis and the decline of the Rust Belt. However, we do not examine these shocks because the farm crisis had little effect on cities and because the Rust Belt was more of a sustained decline than an unanticipated shock.

We examine price changes in cities in the "oil states" listed above and outside the oil states over the period of the oil-price collapse. Since oil prices begin their sharp fall in the first quarter of 1986, we take the previous quarter, the fourth quarter of 1985, as the base period. We choose the first quarter of 1987 as an endpoint because in that quarter employment reaches its trough in Texas, the most populous of the oil states, and because oil prices level off in the first quarter of 1987.[9]

A. Data

For this analysis, we use three types of data: measures of the relative importance of local and national chains at the city level; price data at the city level; and measures of the impact of the oil shock on different cities.

Chain Locations.—The annual publication *Supermarket News Distribution Study of Grocery Store Sales* lists the supermarkets operating in all Metropolitan Statistical Areas (MSA) in the United States as well as the number of stores they operate in each of those

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early model of pricing under imperfection information and its implications for macroeconomics.

[8] Based on the *Quarterly Financial Reports*, the correlation between changes in national income and the ratio of profits to assets in the supermarket industry was 0.06 for the period 1981:1 to 1993:1. Thus, industry profitability is somewhat procyclical; however, it is not as procyclical as the durable goods sector which has a correlation of 0.28.

[9] The LBO of Safeway, an important competitor in the oil states, occurs in 1986:4. However, Chevalier (1995b) suggests that effects of LBOs are not apparent in prices until three to four quarters following an LBO.

TABLE 1—SUMMARY STATISTICS: THE OIL SHOCK OF 1986

Variable	Mean	Variable	Mean
ΔPRICE	−0.006	NATSHARE × OILIMP	0.005
	(0.049)		(0.008)
NATSHARE	0.347	ΔEMP	−0.003
	(0.229)		(0.033)
OILDUM	0.220	NATSHARE × ΔEMP	−0.002
			(0.014)
NATSHARE × OILDUM	0.109	ΔWAGE	0.058
	(0.232)		(0.389)
OILIMP	0.011		
	(0.017)		

Notes: This table reports summary statistics for the sample of 100 observations analyzed in Table 2. ΔPRICE is the percentage change in the local supermarket price index over the period 1985:4 to 1987:1. NATSHARE is the fraction of a city's stores that are owned by national chains. OILIMP is the share of a state's earnings accounted for by oil and gas production. OILDUM is a dummy variable that takes the value 1 if the city is in a state in which OILIMP is greater than 2 percent. ΔEMP is the percentage change in employment in the city's state over this period. ΔWAGE is the percentage change in wages of a sample of workers in sales occupations according to the *Current Population Survey.* Standard deviations are reported in parentheses below the means.

MSAs. We use the 1986 edition of the *Supermarket News* volume which contains data from 1984. From these data we calculate, for each MSA, NATSHARE, the share of stores in a city owned by national chains. We define a national chain as a supermarket chain that operates in more than two of the nine U.S. Census Regions. We determine the number of regions in which a supermarket chain operates using the *Supermarket News* data. As Table 1 indicates, in our sample, on average, national chains comprise 34.7 percent of the stores in a city.

Prices.—Quarterly price data are drawn from the American Chamber of Commerce Researchers Association (ACCRA) *Cost of Living Index*. The ACCRA data are collected in surveys by local chambers of commerce under guidelines set by the American Chamber of Commerce Researchers Association. Included in these data are quarterly prices for 27 specific grocery products (for example, 18 oz. Kellogg's Corn Flakes, a 12 oz. can of Minute Maid frozen orange juice). The recorded price is the average price of a sample of supermarkets in the city.

We use ACCRA's index of grocery prices which is composed of a weighted basket of grocery products. The index for a city is the weighted sum of the price of each item in the ACCRA basket divided by the price of that item in all cities in the ACCRA sample. For each sample period, ACCRA covers approximately 260 cities.

We have data on prices and NATSHARE for a total of 100 MSAs over the time period, including 22 cities in the oil states. We calculate one price index per MSA by averaging the price indices of all observations within the MSA that are available for both 1985:4 and 1987:1. We use the same cities in an MSA in the two quarters so that the price indices are comparable. For the majority of MSAs in the *Supermarket News* data, the only price data available are from the central city of the MSA.

Our measure of the price change, ΔPRICE, is just the percentage change in the price index over the period, that is the price index in 1987:1 less the price index in 1985:4 divided by the price index in 1985:4.

Impact of the Oil Shock.—We use three different measures of a city's exposure to the oil shock. The first is just a dummy variable, OILDUM, which takes the value one if the city is in an oil state. Using data from the Regional Economic Service, we define an oil state as one where the share of total earnings accounted for by oil and gas extraction in 1985,

OILIMP, exceeds 2 percent. Texas, Louisiana, New Mexico, Oklahoma, Colorado, Wyoming and Alaska meet this criterion. The average value of OILIMP for these states is 4.8 percent and the largest is Wyoming with a value of 7.4 percent. In these states employment falls on average by 7.2 percent during the period as compared to a fall of 0.3 percent for the sample as a whole.

The second measure we use is just the continuous measure of a city's exposure to oil industry, OILIMP. The mean value of OILIMP in the sample is 1.1 percent.

The final measure is, ΔEMP, the percent change in employment in the city's state from the peak to the trough. This variable is calculated from quarterly state data published by Data Resources, Inc. We use changes in employment to measure the condition of the local economy rather than unemployment because, as mentioned above, workers tend to leave states that are in a recession. Thus, the unemployment numbers tend to understate the true effects of the recession. In the sample there is a strong negative correlation of -0.83 between ΔEMP and OILIMP, indicating that employment falls more in states with high exposure to the oil industry.

One of the controls we use in the analysis is ΔWAGE, the percentage change in wages of workers in ''sales occupations'' at the MSA level during the period. We construct ΔWAGE by using individual level data from the *Current Population Survey*. This database reports the quarterly wages of a sample of individuals as well as their occupations and the MSA in which they work. For a given quarter in a given MSA, we calculate the average wage of workers in sales occupations as defined in the *Current Population Survey*.[10] The average value of ΔWAGE for the entire sample is 0.058 while it is -0.005 in the oil states.

B. Regression Results

Table 2 reports several regression specifications which test the hypothesis that mark-

ups rose for local chains in the oil states during 1986–1987. In each regression the change in price for a city, ΔPRICE, is regressed on NATSHARE, NATSHARE interacted with various measures of the local economy (OILDUM, OILIMP, ΔEMP), the measure of the local economy itself, and ΔWAGE.

Column (1) lists the results when OILDUM is used to measure the condition of the local economy. The coefficient of NATSHARE is close to zero and statistically insignificant. This variable is included to control for the possibility that the prices of national and local chains were changing at different rates over this period. However, the coefficient suggests that this was not the case.

The coefficient of interest, that of the interaction term NATSHARE $\times$ OILDUM, is -0.115 which is statistically significant at the 2-percent level. The effect of NATSHARE on the price change in the oil states is the sum of the coefficients of NATSHARE and NATSHARE $\times$ OILDUM which is -0.108. This effect is also statistically significant at the 2-percent level. This indicates that in the oil states prices fall more in cities with a large presence of national chains, precisely what the theory predicts.

The estimated effect of NATSHARE on the price change in the oil states is quite large. A one standard deviation increase in NATSHARE from its mean value of 0.35 to 0.58 decreases the expected percentage change in the price index from -0.020 to -0.045 for a city in the oil states. For reference, the standard deviation of percentage price changes for all cities in the ACCRA data base is 0.048.

The regression also includes two other variables as proxies for demand shocks and cost shocks. OILDUM is included because the fall in demand in the oil states could reduce prices in those states relative to other states. This effect would suggest a negative coefficient of OILDUM; however, the coefficient is positive and statistically insignificant. This could be because the upward pressure on prices from the increased markups of local chains in the oil states offsets the downward pressure on prices from the drop in demand.

We also include ΔWAGE as a crude control for the possibility that labor costs rise in cities

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[10] Workers in sales occupations no doubt include those that do not work in supermarkets and exclude some that do, but it is the closest measure we have of the wages of supermarket workers at the local level.

TABLE 2—REGRESSION RESULTS: THE OIL SHOCK OF 1986

Variable	Coefficients		
	(1)	(2)	(3)
NATSHARE	0.007	−0.006	−0.038
	(0.305)	(−0.223)	(−1.852)
NATSHARE × OILDUM	−0.115		
	(−2.270)		
OILDUM	0.017		
	(0.651)		
NATSHARE × OILIMP		−3.033	
		(1.724)	
OILIMP		0.721	
		(0.964)	
NATSHARE × ΔEMP			2.267
			(2.365)
ΔEMP			−0.641
			(−1.570)
ΔWAGE	0.007	0.008	0.007
	(0.59)	(0.641)	(0.565)
Constant	0.000	0.002	0.009
	(0.000)	(0.200)	(0.999)
$\bar{R}^2$	0.170	0.102	0.130

Notes: This table reports regression results in which the dependent variable is ΔPRICE, the percentage change in a city's price index from 1985:4 to 1987:1. NATSHARE is the fraction of a city's stores that are owned by national chains. OILIMP is the share of a state's earnings accounted for by oil and gas production. OILDUM is a dummy variable that takes the value 1 if the city is in a state in which OILIMP is greater than 2 percent. ΔEMP is the percentage change in employment in the city's state over the period. ΔWAGE is the percentage change in wages of a sample of workers in sales occupations according to the *Current Population Survey*. There are 100 observations. *t* statistics are in parentheses below the estimated coefficients.

in oil states that are dominated by local chains. The coefficient of this variable is positive as one might expect but it is statistically insignificant.

Column (2) of Table 2 uses the continuous measure of the importance of oil to the state's economy, OILIMP. We regress the price change, ΔPRICE on NATSHARE, the interaction term NATSHARE × OILIMP, OIL-IMP, and ΔWAGE.

Like the regression in column (1), the coefficient of NATSHARE is indistinguishable from 0, indicating that cities in states where earnings from oil are zero, NATSHARE has no effect on supermarket prices. However, the interaction term, NATSHARE × OILIMP is negative and statistically significant at the 10-percent level: the contraction in demand in cities with large exposures to the oil industry leads to a larger price reduction (smaller price increase) in cities where NATSHARE is high.

That is, local chains are more prone to increase prices relative to national chains in cities that are more adversely affected by the oil shock. Neither OILIMP nor ΔWAGE has a statistically significant coefficient. The estimated coefficients imply that, for a city with OILIMP of 0.028, one standard deviation above the mean, a one standard deviation increase in NATSHARE from its mean of 0.35 to 0.58 increases the expected percentage price change from −0.009 to −0.021.

Column (3) of Table 2 replaces OILIMP with ΔEMP (the percent change in employment) which is a more direct measure of the effect of the shock on the state's economy. As mentioned, there is a strong negative correlation between ΔEMP and OILIMP, indicating that exposure to the oil shock is a good indication that employment will fall.

We regress ΔPRICE on NATSHARE, the interaction term NATSHARE × ΔEMP, ΔEMP,

and ΔWAGE. Consistent with our other results, we find that the coefficient of NAT-SHARE is statistically indistinguishable from 0. The coefficient of NATSHARE $\times$ ΔEMP is positive and statistically significant; as before, in cities in which demand contracts, prices fall more if NATSHARE is large.[11]

The coefficient implies that, for a city with employment growth of -0.033, one standard deviation below the mean, a one standard deviation increase in NATSHARE from its mean of 0.35 to 0.58 decreases the expected percentage price change from -0.009 to -0.035.

IV. LBOs and the 1990–1991 Recession

In this section we examine the effect of the macroeconomic recession of 1990–1991 on the pricing behavior of supermarket chains. Many of these chains undertook leveraged buyouts during the latter half of the 1980's. One of the main effects of an LBO—indeed, one of the functions of an LBO, according to Michael Jensen (1986)—is to limit the cash available to corporate managers by committing them to make large principal and interest payments. As a result, the recession is likely to have put greater pressure on LBO firms to boost short-run cash flows so as to be able to make their higher principal and interest payments. Thus, we would expect LBO firms to have raised prices relative to non-LBO firms particularly in cities hit hardest by the recession.

A. Data

The peak of the cycle according to the National Bureau of Economic Research was July 1990 and the trough was March 1991. As with the other episodes, we calculate the price changes from one quarter before the peak to the quarter of the trough. Thus, we examine

price changes over the period 1990:2 to 1991:1.[12] The price data, as before, are from ACCRA.

For information about supermarket locations, we use the 1992 edition of Progressive Grocer's *Market Scope*, which contains data for 1991. *Market Scope* lists all of the supermarkets operating in each of the 100 largest MSAs in the United States and lists the number of stores in each chain. However, *Market Scope* lists store names, not the names of parent companies so store names were matched to parent company names using the *Retail Tenants Directory, Thomas's Grocery Register*, and supermarket firms' annual 10-K reports.[13] *Market Scope* only covers the 100 largest MSAs.

We identify supermarket chains as having undertaken an LBO by first searching for listings of supermarket LBOs in quarterly editions of *Mergers and Acquisitions*, which contains information on all ownership transactions (including LBOs) of greater than $1 million. Second, we search all references to transactions involving the supermarket parent companies in the sample using indices to *Supermarket News, Supermarket Business*, and *Progressive Grocer*. All of the LBOs examined in this study were completed between 1981 and 1990.

For each of the cities in *Market Scope* we calculate LBOSHARE, the share of all stores in a city owned by firms which undertook LBOs prior to the second quarter of 1990. As Table 3 indicates the mean of the variable in our sample is 24.6 percent. We then match these data with local price data collected from ACCRA. We calculate one price index per

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[11] If one is concerned about ΔEMP as a proxy for the demand shock one can use OILIMP to instrument ΔEMP and NATSHARE $\times$ OILIMP to instrument NATSHARE $\times$ ΔEMP which we did. The coefficient estimates were very similar to those reported in the table and the coefficient estimate of NATSHARE $\times$ ΔEMP was statistically significant at the 10-percent level.

[12] Some have argued that the recession should really be dated from 1990:2 to 1991:4 because the recovery in 1991:1 and 1991:2 was minimal. The results presented here are very similar for that time period.

[13] The Progressive Grocer data were used in this part of the study rather than the Supermarket News data because it is higher quality data. Supermarket News, however, publishes data for all MSAs, while Progressive Grocer has data for only the 100 largest MSAs. We use the Supermarket News data for the oil episode for two reasons: we needed to obtain data for as many cities in the oil producing states as was possible; and we needed to measure to what extent the supermarket chains were nationwide.

TABLE 3—SUMMARY STATISTICS: 1990–1991 RECESSION

Variable	Mean
ΔPRICE	0.006
	(0.029)
LBOSHARE	0.246
	(0.194)
ΔEMP	−0.022
	(0.016)
LBOSHARE $\times$ ΔEMP	−0.005
	(0.006)
ΔWAGE	0.035
	(0.236)

Notes: This table reports summary statistics for the sample of 59 observations analyzed in Table 4. ΔPRICE, the percentage change in a city's price index from 1990:2 to 1991:1. LBOSHARE is the fraction of a city's stores that are owned by chains that undertook a leveraged buyout during the 1980's. ΔEMP is the percentage change in employment in the city's state during the period. ΔWAGE is the percentage change in wages of a sample of workers in sales occupations according to the *Current Population Survey*. Standard deviations are reported in parentheses below the means.

TABLE 4—REGRESSION RESULTS: 1990–1991 RECESSION

Variable	Coefficient
LBOSHARE	−0.024
	(0.788)
LBOSHARE $\times$ ΔEMP	−1.970
	(1.767)
ΔEMP	0.384
	(1.005)
ΔWAGE	0.0173
	(1.052)
Constant	0.010
	(0.864)
$\bar{R}^2$	0.095

Notes: This table reports regression results where the dependent variable is ΔPRICE, the percentage change in a city's price index from 1990:2 to 1991:1. LBOSHARE is the fraction of a city's stores that are owned by chains that undertook a leveraged buyout during the 1980's. ΔEMP is the percentage change in employment in the city's state during the period. ΔWAGE is the percentage change in wages of a sample of workers in sales occupations according to the *Current Population Survey*. There are 59 observations. *t* statistics are in parentheses below the estimated coefficients.

MSA by averaging the price indices of all observations within the MSA that are available for both 1990:2 and 1991:1.

We use the same set of cities in both quarters to construct a comparable index. For most MSAs in *Market Scope* the only price data are from the central city of the MSA. There are 59 observations for which we have price data from ACCRA and store location data from *Market Scope*. For reference, the means and standard deviations of the variables are listed in Table 3.

B. *Regression Results*

We regress the percentage change in the price index from 1990:2 to 1991:1, ΔPRICE, on LBOSHARE, an interaction term, LBOSHARE $\times$ ΔEMP, ΔEMP, and ΔWAGE. The variables ΔEMP and ΔWAGE are defined over the period 1990:2 to 1991:1. The results are presented in Table 4.

While LBOSHARE has no statistically significant effect on ΔPRICE, the interaction term is negative and statistically significant at the 8-percent level, which is consistent with the theory. The coefficients of both ΔEMP

and ΔWAGE are positive, but statistically insignificant.

In a city in which employment growth is −0.5 percent (one standard deviation above its mean), then even a one standard deviation increase in the share of LBO firms from its mean of 0.25 to 0.44 actually lowers the expected price increase in the city slightly from 0.5 percent to 0.2 percent. However, in a city very adversely affected by the recession, an increase in the share of LBO firms leads to higher prices. In a city with employment growth of −3.8 percent, (one standard deviation *below* its mean), a one standard deviation increase in the LBOSHARE from its mean of 0.25 to 0.44 more than doubles the expected price increase, from 0.8 percent to 1.8 percent. Thus, the effects are both statistically and economically significant.

V. LBOs Following the 1990–1991 Recession: Firm-Specific Data

The theory outlined in Section I has implications for firm-level prices in addition to implications for market-level prices analyzed above. First, the theory predicts that a firm's

price will be more countercyclical if it is more liquidity constrained. And, second it predicts that a firm's price will be more countercyclical if its *rivals* are more liquidity constrained; this follows from the fact that prices are strategic complements. In this section, we use local firm-level data on supermarket pricing to analyze the effect of liquidity constraints on a firm's pricing as well as its rivals' pricing.

The firm-level data we use begins in the first quarter of 1991. According to the NBER, this is the end of the 1990–1991 recession. However, the recovery from this recession was weak for several quarters following the trough. In fact, many states in the northeast and California experienced negative growth in employment until the fourth quarter of 1992. Thus, even though our price data do not coincide with the macroeconomic recession, we can use the geographic heterogeneity in the recovery to investigate the impact of liquidity constraints on the pricing behavior of LBO and non-LBO firms.

A. Data

The data on supermarket prices are from Information Resources, Inc. IRI collects these data from electronic product scanners. The IRI data are firm specific, but average together prices for all of a firm's stores within the market area defined by IRI. An IRI market area is generally somewhat larger than an MSA, and in some cases, an IRI market area covers an entire state.

The *Market Scope* data described in Section IV also provides information on store locations for IRI market areas. We used the 1992 edition of *Market Scope*, which contains data for 1991, to measure the number of stores of different types in each IRI market.

In the analysis, we use a price index that is a quarterly average of prices for a basket of product types.[14] A product type is a category

such as dry pasta, ready-to-eat cereals, diet soft drinks, and so on. The price of a product type observed for a supermarket chain in an IRI market area for a quarter is the average price of a unit of the good, averaged across all sales in the quarter. For example, prices for boxes of cereal of different brands and sizes are recorded in dollars per pound and averaged over the quarter. Out of the 50 product types with the highest dollar sales, 38 were scanned by all supermarkets in the IRI sample. We use these 38 product types for our sample.

We examine price changes over two periods: the period from 1991:1 to 1991:4 and the period from 1991:1 to 1992:4.[15] We have 110 observations of firms in local markets for the shorter time period. We have a smaller set of firm-market pairs with complete data for the longer time period (89 observations).

To test the prediction that cash-constrained firms charge higher prices, we construct LBO, a dummy variable which equals 1 if the firm is an LBO firm. As Table 5 indicates, for the larger sample of 110 firms, 38.1 percent of the firms are LBO firms. We use OLBOSHARE to measure the second effect, whether firms charge higher prices when their rivals are cash constrained. For each firm in each market, we calculate OLBOSHARE as the share of stores in that local market owned by LBO chains other than the firm itself. The mean value of OLBOSHARE is 14.9 percent.

As our measure of the condition of the local economy, we use the percent change in employment in the state of the market area, ΔEMP. In some cases, the IRI market areas include counties from several states. In these cases, we construct ΔEMP as a weighted average of the percent change in employment in each of the states included in the market area. The weights are the share of the population in the IRI market area accounted for by residents of each state. The average value of ΔEMP is roughly 2 percent which is consistent with the view that the economy is in recovery. But in some states, such as California, employment growth is negative.

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[14] These prices do not include deductions for manufacturers' or store coupons. We suspect that the bias this introduces is small. For coupons to cause an important bias, there would have to be a systematic difference in the extent to which LBO versus non-LBO *change* their prices via coupons over the business cycle.

[15] Note that unlike the previous parts of the study these are nominal price changes but since we are comparing price changes across firms this is of no consequence.

TABLE 5—SUMMARY STATISTICS: 1990–1991 RECESSION USING FIRM-SPECIFIC DATA

Variables	Mean 1991:1–1991:4	Mean 1991:1–1992:4
ΔPRICE	0.023	0.080
	(0.028)	(0.031)
LBO	0.382	0.393
LBO × ΔEMP	0.007	0.010
	(0.012)	(0.021)
OLBOSHARE	0.149	0.151
	(0.151)	(0.146)
OLBOSHARE × ΔEMP	0.002	0.004
	(0.004)	(0.007)
ΔEMP	0.020	0.030
	(0.015)	(0.028)
ΔWAGE	0.053	0.124
	(0.221)	(0.266)

Notes: This table reports summary statistics for the sample analyzed in Table 6. ΔPRICE is the percentage change in a firm's price index for a particular city in one of two periods, 1991:1 to 1991:4 and 1991:1 to 1992:4. LBO is a dummy variable which takes the value of one if the firm had previously undertaken a leveraged buyout. ΔEMP is the percentage change in employment in the city's state during the period. OLBOSHARE is the share of stores in the local market owned by LBO chains other than the firm itself. ΔWAGE is the percentage change in wages of a sample of workers in sales occupations according to the *Current Population Survey.* There are 110 observations in the shorter sample and 89 observations in the longer sample. Standard deviations are reported in parentheses below the means.

B. Regression Results

We regress the percent change in the price of the food basket in each of the 110 firm-market pairs from 1991:1 to 1991:4 on LBO, OLBOSHARE, an interaction term LBO × ΔEMP, an interaction term OLBOSHARE × ΔEMP, ΔEMP, and ΔWAGE. The coefficients are estimated using OLS with OLS standard errors. Because there are multiple firm observations for a single city and multiple city observations for a single firm, OLS standard errors could be biased due to a correlation of the disturbance term across related observations. Brent R. Moulton (1986) proposes a version of the T. R. Breusch and A. R. Pagan (1980) Lagrange multiplier method to test for the appropriateness of OLS standard errors in contexts very similar to ours. Following Breusch and Pagan, we extend the test to allow for the possibility of two types of error correlation: intrafirm and intracity. We fail to reject the null hypothesis that there is no intracity or intrafirm correlation at the 25-percent confidence level.

Therefore OLS estimates and standard errors are not subject to this potential problem.

The results for this regression are reported in the first column of Table 6. The regression shows that LBO firms tend to raise prices more than non-LBO firms; the coefficient of LBO is positive and statistically significant. This could reflect marginal cost increases of LBO firms relative to non-LBO firms and need not be associated with relative increases in the markup. However, the coefficient of the interaction term LBO × ΔEMP is negative and statistically significant. Thus, LBO firms tend to raise price more relative to non-LBO firms in local markets in which economic conditions are worse. This is not consistent with the view that there is a relative increase in the marginal costs of LBO firms, but rather that markups of LBO firms rise relative to non-LBO firms.

In a "boom" city, with employment growth of 3.5 percent (one standard deviation above the mean), a non-LBO firm increases prices by more than a LBO firm, 2.3 percent versus 1.4. However, in the typical "bust" city, LBO firms raise prices more than non-LBO firms.

TABLE 6—REGRESSION RESULTS: 1990–1991 RECESSION USING FIRM-SPECIFIC DATA

Variables	Coefficients	
	1991:1–1991:4	1991:1–1992:4
LBO	0.030	0.025
	(3.235)	(2.670)
LBO × ΔEMP	−1.131	−0.402
	(−2.942)	(−1.659)
OLBOSHARE	0.109	0.120
	(3.454)	(3.456)
OLBOSHARE × ΔEMP	−2.834	−1.152
	(−2.184)	(−1.389)
ΔEMP	0.712	0.375
	(2.578)	(1.807)
ΔWAGE	−0.004	−0.011
	(−0.333)	(−0.892)
Constant	−0.003	0.050
	(−0.413)	(5.044)
$\bar{R}^2$	0.179	0.200
Number of observations	110	89

Notes: This table reports regression results where the dependent variable is ΔPRICE, the percentage change in a firm's price index for a particular city over two different time periods, 1991:1, 1991:4 and 1991:1–1992:4. ΔEMP is the state employment change during the period; LBO is a dummy variable equal to 1 if the firm undertook an LBO; LBOEMP is LBO multiplied by ΔEMP; OLBOSHARE is a measure of the market shares of the other supermarket chains that undertook leveraged buyouts; OLBOEMP is OLBOSHARE multiplied by ΔEMP. There are 110 observations in the shorter sample and 89 observations in the longer sample. *t* statistics are in parentheses below the estimated coefficients.

With employment growth one standard deviation below the mean (a city with 0.5 percent employment growth), an LBO firm increases prices by 3.8 percent while a non-LBO firm increases prices by 1.4 percent.

If LBO firms increase their prices by 2.4 percent more than non-LBO firms, this can have an economically significant effect on short-run profits. One can show that for each percentage increase in prices, pretax net profits rise by $(1 - g\varepsilon)/n$ percent where g is the gross profit margin, n is the net profit margin, and ε is the short-run elasticity of demand. In the *Value Line* sample of supermarket chains, the average pretax net profit margin is roughly 2.5 percent and the gross profit margin is roughly 25 percent. With a short-run elasticity of demand of 0.3 as estimated by Hoch et al. (1993), each 1-percent price increase would raise net profits by 37 percent. Thus, a 2.4-percent price increase would raise profits by about 89 percent which is obviously quite large. Of course, the fall in demand during the recession tends to lower profits. But, it is clear

that seemingly small price increases can offset the effects of a demand drop.[16]

This regression also shows that firms tend to raise prices more when their rivals are highly leveraged. The coefficient of OLBO-SHARE is positive and statistically significant. This is evidence that leverage leads rivals to increase prices (for rivals) and that prices are strategic complements. In addition, the coefficient of the interaction term OLBOSHARE × ΔEMP is negative and statistically significant. Thus, firms tend to raise price more when their rivals are leveraged and they are competing in a local market that is experiencing slow economic growth.

We consider first a "boom" city with employment growth of 3.5 percent (one standard

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[16] In fact, one can show that if marginal costs do not change as a result of the recession, each 1-percent drop in demand lowers net profits by 10 percent (that is, g/n percent). Thus, a 2.4-percent price increase would leave profits unchanged when demand drops 8.9 percent.

deviation above the mean). An increase in the share of other LBO firms from its mean of 14.9 percent to 30 percent (one standard deviation above the mean) increases the price trivially from 2.3 percent to 2.5 percent. The effect is even larger in markets with low employment growth; with an employment growth of 0.5 percent, an increase in OLBOSHARE by one standard deviation from 14.9 percent to 30.0 percent would more than double the non-LBO firm's price increase, from 1.4 percent to 2.9 percent.

The second column of Table 6 repeats this specification for the longer time period, 1991:1 to 1992:4. The results are qualitatively similar, but the coefficient of OLBOSHARE $\times$ ΔEMP is statistically significant only at the 17-percent level. The rest of the coefficients continue to be statistically significant at greater than the 10-percent level.[17]

VI. Conclusion

The theory and evidence presented here suggest that capital-market imperfections induce liquidity-constrained firms to increase markups during recessions and lower them during booms. This, in turn, may amplify the effects of demand shocks on output. One might argue, however, that capital-market imperfections are not likely to have an effect on the macroeconomy because firms which are not liquidity constrained will take up the slack. Our model and evidence points to the limitations of this argument. Since prices are strategic complements, unconstrained firms increase their markups during recessions in response to their constrained rivals. Thus, in essence, even though some firms are not liquidity constrained, they act as if they are constrained. This suggests that shocks to one set of firms—rather than being stabilized by

other firms in the industry—are transmitted to other firms in the industry.

Of course, it may be difficult to apply our results to the economy as a whole because we study only one industry. It remains to be seen whether liquidity constraints can explain countercyclical markups for the macroeconomy. But, our evidence of countercyclical markups in an industry that is not particularly cyclical suggests that the effects may be even more important in other industries.

In fact, in Chevalier and Scharfstein (1995) we reexamined the cross-industry study of Rotemberg and Woodford (1991) to see whether liquidity constraints could explain countercyclical markups in a range of manufacturing industries. Rotemberg and Woodford found that markups are more countercyclical in more concentrated industries which they took to be evidence of countercyclical collusion. However, this finding is also consistent with the view that markups are countercyclical because the incentive to price for market share is procyclical since in unconcentrated, competitive industries there is no incentive to price for market share. To see whether liquidity constraints could help explain the results we added to their basic regression a measure of the market share of small firms, those with assets under $25 million. These firms are more likely to be liquidity constrained during a downturn. We found that, controlling for concentration, markups are more countercyclical in industries where small firms are more important, consistent with our theory. This finding is, however, inconsistent with Rotemberg and Woodford's interpretation of countercyclical collusion. In their model, liquidity-constrained small firms have more incentive to cut prices during recessions because in a collusive oligopoly it increases short-run profits.

Our results also point to another way in which shocks can be transmitted through the economy. We find that firms with high cash flows in divisions outside the oil states were better able to invest in market share during the recession in the oil states. By analogy, one would expect firms with low cash-flow divisions to invest less in their other divisions. Indeed, Owen Lamont (1996) finds that the non-oil divisions of oil companies cut investment following the large drop in oil prices in

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[17] The basic results in Table 6 are robust to the inclusion of other controlling variables, such as the store share of the firm, the store share of the firm interacted with the LBO dummy, the store share of the firm interacted with the employment change, and the store share of the firm interacted with both the LBO dummy and the employment change. None of these other variables had statistically significant coefficients in either time period, however.

1986. Thus, a shock to one sector of the economy was transmitted to other sectors of the economy through the internal capital allocation mechanisms of corporations. Of course, the higher cash flows of the non-oil divisions may have reduced the effects of the shock to the oil sector. Whether, on net, conglomerates tend to transmit shocks or stabilize them is an empirical issue.

Finally, we note that this paper fits into a large body of work indicating that liquidity constraints can also have real effects on capital investment, inventory investment, and employment (see Bernanke et al. [1993] for a review). The studies tend to look at each of these factors separately, without attempting to relate the findings to each other. However, one would expect, for example, that if firms cut back on inventory investment during recessions because of liquidity constraints, they would also increase markups for the same reason. This suggests that one should look not just at the movements of inventories, but also look at the comovements of inventories with markups. This is a standard approach used to test other macroeconomic theories and it should be applied to these theories as well.

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