Flexible Prices and Leverage*

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Abstract

The frequency with which firms adjust output prices is an important determinant of persistent differences in capital structure across firms. We show the most flexible-price firms have a 19% higher long-term financial leverage than the most inflexible-price firms, controlling for known determinants of capital structure. We rationalize this novel fact in a costly-state-verification model, in which inflexible-price firms are more exposed to aggregate shocks, and face tighter financial constraints. In the model, bank lending relaxes financial constraints through monitoring and narrows the gap in financial leverage between inflexible- and flexible-price firms. Consistently, inflexible-price firms increased leverage more than flexible-price firms following the staggered implementation of the Interstate Bank Branching Efficiency Act across states and over time, which we use in a triple-differences identification strategy. Firms' frequency of price adjustment did not change around the deregulation.

JEL classification: E12, E44, G28, G32, G33

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I Introduction

Understanding firms' capital structure is perhaps the central question in corporate finance. This paper argues the frequency with which firms adjust product prices can help us understand persistent differences in financial leverage across firms (Lemmon, Roberts, and Zender (2008), DeAngelo and Roll (2015)). We test whether firms' inability to adjust output prices to aggregate and idiosyncratic shocks exposes them to financial constraints, and hence reduces leverage.

Price rigidity—the fact that firms do not adjust prices to macroeconomic shocks—has long been a focus in Macroeconomics and Industrial Organization. We build on the evidence of persistent heterogeneity in the firm-level frequency of price adjustment across and within industries (Nakamura and Steinsson (2008) and Gorodnichenko and Weber (2015)). Golosov and Lucas (2007) and Alvarez et al. (2011) show firms' frequency of price adjustment changes little over time, even with inflation rates ranging from 0% to 16%. Weber (2015) documents the correlation between price rigidity and the cross section of stock returns. Firm-level price rigidity is highly persistent and a source of systematic risk, making it a viable candidate to explain the capital structure of firms.

To guide our empirical analysis, we develop a costly-state-verification model, in which firms adjust prices imperfectly to macroeconomic shocks. Firms with inflexible prices are more exposed to aggregate shocks, and hence their profits are more volatile. Our contribution is proposing price inflexibility as an economically motivated driver of profit volatility. Because of higher profit volatility, inflexible-price firms have an incentive to under-report their realized profits when monitoring is costly, and can pledge credibly a lower share of their expected profits ex ante to obtain financing compared to flexible-price firms. Therefore, inflexible-price firms face more binding financial constraints. In the model, bank lending relaxes financial constraints by providing a monitoring technology.

The model predicts inflexible-price firms have unconditionally lower financial leverage than flexible-price firms. Moreover, a positive shock to the quality of bank monitoring increases the leverage of inflexible-price firms more than the leverage of flexible-price firms. Empirically, uncertainty about the price level and the role of price-setting frictions is most relevant for profits over long horizons. In our baseline analysis, we therefore test these predictions using long-term leverage as the main outcome variable, but all our results go through if we look at the short-term leverage or the total leverage of firms.

We first document a novel stylized fact consistent with the model's predictions, that is, flexible-price firms have higher financial leverage than inflexible-price firms. We document this fact using the confidential micro data underlying the official producer price index (PPI) of the Bureau of Labor Statistics (BLS). We observe monthly good-level pricing data for a subsample of the S&P500 firms from January 1982 to December 2014.

Our preferred leverage measure—the long-term debt-to-assets ratio—is 4 percentage points higher for the 25% most flexible-price firms than for the 25% most inflexible-price firms, which is 19% of the average ratio in the sample. A two-standard-deviation increase in our continuous measure of price flexibility is associated with 2.5-percentage-point higher long-term debt-to-assets ratio, which is 12% of the average ratio in the sample. We estimate these magnitudes after partialling out known determinants of capital structure such as size, tangibility, profitability, and the book-to-market ratio. We also control for industry concentration and for firm-level measures of market power, which might be correlated with the degree of price flexibility of firms. The results are similar if we only exploit the variation in price flexibility within industries and within years. This result is important, because product market considerations at the industry level affect firms' demand for debt (e.g., see Maksimovic (1988) and Maksimovic (1990)). The size of the estimated coefficients does not change when we use the errors-in-variables estimator based on linear cumulant equations of Erickson, Jiang, and Whited (2014).

A crucial feature of our model is that inflexible-price firms are riskier borrowers, because their profits are more volatile. We show firms with inflexible prices are more likely to default compared to flexible-price firms. This fact adds to previous findings that firms with inflexible product prices have more volatile cash flows after monetary policy shocks, and have higher total and idiosyncratic stock return volatility (Gorodnichenko and Weber (2015) and Weber (2015)).

To assess whether the effect of price flexibility on leverage is causal, one route would be testing the effect of a shock to firm-level price flexibility on leverage, or proposing an instrument for price flexibility. But price flexibility is a highly persistent characteristic of firms. For instance, in our sample, a firm-level regression of post-1996 price flexibility onto pre-1996 price flexibility yields a slope coefficient of 93%, and we fail to reject the null that the coefficient equals 1 at any plausible level of significance.[1] This persistence

[1]See also Nakamura and Steinsson (2008), Golosov and Lucas (2007), and Alvarez et al. (2011).

suggests we can hardly consider a shock to firm-level price flexibility for identification purposes in our sample.

We therefore propose an identification strategy derived from our model's predictions, and in line with the financial constraints literature. We (i) identify a positive shock to the quality of monitoring of the banks firms can access, (ii) show inflexible-price firms increase leverage more than flexible-price firms after the shock, and (iii) show the effect does not revert in the short run. Our strategy exploits a quasi-exogenous shock to financial constraints, and uses ex-ante unconstrained firms to assess the causal effect of financial constraints on inflexible-price firms.

The shock we use is the staggered state-level implementation of the Interstate Banking and Branching Efficiency Act (IBBEA) between 1994 and 2005 (Rice and Strahan (2010) and Favara and Imbs (2015). Restrictions on U.S. banks' geographic expansion date back at least to the 1927 McFadden Act, but the IBBEA of 1994 let bank holding companies enter other states and operate branches across state lines. The step-wise repeal of interstate bank branching restrictions increased the supply of credit. Banking deregulation results in lower interest rates charged (Jayaratne and Strahan (1996)), more efficient screening of borrowers (Dick and Lehnert (2010)), increased spatial diversification of borrowers (Goetz, Laeven, and Levine (2013)), higher loan volume (Amore, Schneider, and Žaldokas (2013)), and more credit cards (Kozak and Sosyura (2015)).

We interpret the staggered state-level implementation of the IBBEA as a positive shock to banks' monitoring effectiveness, which is exogenous to the individual firms' financial decisions. We therefore propose a triple-differences identification strategy. We compare outcomes within firms before and after the implementation of the IBBEA in the state where the firms are headquartered, across firms in states that deregulated or not, and across flexible- and inflexible-price firms. Firms in states that did not deregulate act as counterfactuals for the evolution of the long-term debt of treated firms absent the shock.

To assess the plausibility of the identifying assumptions our strategy implies, we show that before the shock, the trends of long-term debt of flexible- and inflexible-price firms are parallel, and the price flexibility of firms does not change around the shock.

Consistent with our model's predictions, inflexible-price firms increased their leverage more than flexible-price firms after the deregulation. Crucially, the effect is driven by inflexible-price firms with a lower cash-to-assets ratio, which were more likely to need external financing to fund their operations. The most flexible-price firms kept their leverage virtually unchanged after the deregulation. Note our results do not vary across firms that have an investment grade or not, which is important to verify because Faulkender and Petersen (2006) show firms that access the public bond markets have higher leverage.

The availability of product-pricing micro-data requires that we focus on large firms. That these large firms depend on bank credit might appear surprising, but using data from Sufi (2009), we find 95% of the firms in our sample have credit lines with at least one bank. The average rate of utilization is above 20%, which suggests bank relationships are relevant in our sample. These facts are consistent with Beck, Demirguc-Kunt, and Maksimovic (2008), who find large firms are more likely than small firms to rely on bank finance. Moreover, both the likelihood of having credit lines and their sizes increase after the implementation of the IBBEA. Consistent with our results on leverage, inflexible-price firms drive the increase in the size of credit lines.

We assess the validity of our triple-differences results with two falsification tests. We split states into early deregulators (between 1996 and 1998) and late deregulators (after 2000). We first replicate our results comparing flexible- and inflexible-price firms in early and late deregulating states, before and after 1996, but only using observations prior to 2000, when late deregulators had not yet implemented the IBBEA. The effects of the deregulation on capital structure are driven by inflexible-price firms in early deregulation states, as opposed to inflexible-price firms in late deregulation states and flexible-price firms in all states.

In the first falsification test, we repeat the analysis using only observations prior to 1996, when *no state* had yet deregulated. We use 1992 as a placebo implementation date for early deregulators, which is four years before 1996. We do not find any difference in the capital structure of inflexible-price firms in early states compared to inflexible-price firms in late states before and after 1992.

In the second falsification test, we repeat the analysis using observations prior to 1996 *and* after 2000. Before 1996, no states had yet deregulated, and after 2000, all states had deregulated. Consistent with our interpretation of the shock and our model's predictions, inflexible-price firms in both early states and late states have higher long-term debt after

2000 compared to before 1996, whereas flexible-price firms in both sets of states do not change their capital structure after 2000.

Our paper adds to a recent literature studying the macroeconomic determinants of financial leverage, default risk, and bond yields. Bhamra, Kuehn, and Strebulaev (2010) study the effect of time-varying macroeconomic conditions on firms' optimal capital structure choice. Kang and Pflueger (2015) show that fear of debt deflation is an important driver of corporate bond yields. Favilukis, Lin, and Zhao (2015) document that firms with higher wage rigidities have higher credit risk, whereas Simintzi, Vig, and Volpin (2015) show that firms in industries with higher wage rigidities have lower leverage. Labor market frictions vary at the industry level, and hence can hardly account for our findings.

The paper also speaks to the theoretical and empirical literatures that study the effect of volume flexibility on firm's capital structure. The literature is inconclusive about the sign of the effect of volume flexibility on financial leverage. On the empirical side, MacKay (2003) finds volume flexibility reduces financial leverage, whereas Reinartz and Schmid (2015) find the opposite using direct measures of volume flexibility for firms in the utilities sector. On the theoretical side, volume flexibility can decrease default risk (e.g., see Mauer and Triantis (1994)) and promote risk shifting and asset substitution (e.g., see Mello and Parsons (1992)), which have opposite effects on financial leverage in equilibrium. In this paper, we focus on the persistence of capital structure differences across firms, which non-price real flexibility cannot explain. In our empirical analysis, we control for firms' price-to-cost margin, which we define as a linear transformation of operating leverage, to average out the effects of time-varying operating leverage on financial leverage.

II Theoretical Framework

We consider the optimal financing decision of a firm in a one-period partial equilibrium setup with costly state verification (Townsend (1979), Gale and Hellwig (1985)). This stylized model allows us to compare two financing environments. First, the firm borrows through the public bond market. Second, the firm borrows from a bank. Owners of diffusely-owned public bonds might suffer a coordination problem when monitoring private information (Diamond (1991a), Diamond (1991b)). Banks have access to a costly monitoring technology, which distinguishes them from the public bond market.

The model generates two main predictions. First, inflexible-price firms have lower leverage than flexible-price firms. Second, inflexible-price firms increase leverage more than flexible-price firms in response to an increase in monitoring effectiveness.

In the model, firms differ in their ability to adjust output prices to macroeconomic shocks. Inflexible-price firms have greater uncertainty about profits. Their profits are identical to those of flexible-price firms when realized inflation coincides with expected inflation. However, inflexible-price firms have lower profits when realized inflation is either unexpectedly high or unexpectedly low.

Inflexible-price firms have an incentive to report low profits even when profits are high, which limits their debt capacity. Monitoring reduces the incentive to misreport profits, and allows inflexible-price firms to credibly pledge a greater share of real profits to lenders. Bank lending can therefore mitigate the credit constraints inflexible-price firms face.

A. Production and Prices

We use capital letters to denote levels, and small letters to denote logs. The firm's actual price level may differ from the optimal price if the firm can update prices or information only infrequently (Calvo (1983), Mankiw and Reis (2002)). We denote the log difference between actual and optimal product prices by Δp.

For simplicity, the price gap can take three values with associated probabilities

$$Prob\,(\Delta p = 0) = \pi_0, \tag{1}$$

$$Prob\,(\Delta p = h) = \frac{\pi_h}{2}, \tag{2}$$

$$Prob\,(\Delta p = -h) = \frac{\pi_h}{2}, \tag{3}$$

$$\pi_0 + \pi_h = 1. \tag{4}$$

The expected price gap is zero. The parameter h captures how far the firm allows prices to deviate from the optimum when shocks occur to the aggregate price level. The parameter h is a reduced form to model pricing frictions that might originate from costs of price adjustment, managerial costs, information-processing costs, or negotiation costs. Zbaracki et al. (2004) shows a U.S. manufacturing firm with annual revenues of more than \$1bn spends about 1.2% of annual revenues on price adjustments, which corresponds to about 20% of the net profit margin. Gorodnichenko and Weber (2015) calibrate their fully dynamic model to the micro-data underlying the PPI and find similar costs of price adjustments.

In New Keynesian models with partially monopolistic competition, price dispersion leads to production misallocations and real economic costs (Woodford (2003)). When the price gap is negative, firm revenue per unit sold and total firm profits are below the optimum. When the price gap is positive, high prices reduce demand, and firm profits are also below the optimum.

We capture the key features of costly price dispersion with a simple, quadratic profit function. The profit function is maximized at $\Delta p = 0$, ensuring the existence of a flexible-price equilibrium in which all firms charge the same price. Firm profits scale with firm capital K, giving firm profits as

$$Profit_{\Delta p} = K \times R_{\Delta p}, \tag{5}$$

$$R_{\Delta p} = exp\left(r_{\Delta p}\right), \tag{6}$$

$$r_{\Delta p} = \bar{r} - a(\Delta p)^2. \tag{7}$$

Here, $\bar{r} > 0$ and $a > 0$ are constants, reflecting log returns when the price gap is zero and the curvature of the profit function. $\bar{r} > 0$ is needed to ensure a positive net present value return on capital. Equation (5) obtains as a second-order approximation around the optimal price in a micro-founded model. The model predictions do not rely on the specific functional form (5) through (7). We rely on a quadratic profit function to maximize clarity of exposition.

B. The Financing Problem

The owner of the firm has personal wealth or equity, E, which determines the scale of the firm. The owner has all bargaining power, and the lender breaks even in expectation. The interest rate in the economy is zero, and the owner and the investor are both risk neutral. The total capital of the firm is the sum of debt and equity,

$$K = D + E. \tag{8}$$

We make two additional assumptions to make the financing problem interesting. First, we assume the project's net present value is positive; that is,

$$\pi_0 R_0 + \pi_h R_h > 1. \tag{9}$$

Here, $R_0 = exp(r_0)$ and $R_h = exp(r_h)$. Second, we assume the firm's returns are less than 1 in the low-profit state,

$$R_h < 1. \tag{10}$$

Lenders cannot observe firm profits. This assumption captures the idea that lenders cannot costlessly observe firms' optimal and actual pricing strategies. The manager's incentive to misreport realized profits constrains the set of feasible financing contracts. Contracts in our model are real to focus on the cross-sectional implications of the model. With nominal contracts, uncertainty about the aggregate price level can further lower the debt capacity of both inflexible- and flexible-price firms (Fisher (1933), Bhamra, Fisher, and Kuehn (2011), Kang and Pflueger (2015)).

C. Solution without Monitoring

First, we consider the optimal debt contract when no monitoring technology is available. We can think of this setup as a firm that can only borrow from public debt markets.

The optimal contract must satisfy the revelation principle: the borrower reveals her profits truthfully. Without monitoring technology, the optimal financing contract requires constant payments across states. Otherwise, the borrower has an incentive to lie about profits. The project has a positive net present value, and the manager optimally borrows the maximum amount the lender is willing to lend. Optimal leverage follows from the lender's break-even constraint:

$$\frac{D}{K} = R_h. \tag{11}$$

Firms with more inflexible prices, that is, larger h, have lower returns R_h and hence lower leverage.

D. Solution with Monitoring

Next, we consider the case in which the lender can access a costly monitoring technology. This setup resembles a firm that borrows from a bank, which has a costly technology to monitor the manager's activities.

Monitoring costs are proportional to firm size, and are given by γK. Monitoring larger firms requires more effort than monitoring smaller firms. When monitoring is unsuccessful, which occurs with probability $1-\rho$, the lender acquires no information about firm profits. When monitoring is successful, the lender observes the true level of profits, and contract payoffs can be contingent on the monitoring result. We assume monitoring costs are small relative to the expected gains from monitoring:

$$\rho\left(\pi_0 R_0 + \pi_h R_h - 1\right) \quad > \quad \pi_h \gamma. \tag{12}$$

Let C_0 denote the manager's consumption in state 0. The optimal contract gives the manager zero consumption in state h and when he is caught misreporting profits, thereby minimizing the incentives to misreport firm profits in the high-profit state.

The optimal contract maximizes the manager's expected consumption,

$$V \quad = \quad \pi_0 C_0, \tag{13}$$

subject to the following incentive-compatibility constraints:

$$\begin{aligned} C_0 \quad &\geq \quad (1-\rho)K(R_0 - R_h), &(14)\\ C_0 \quad &\leq \quad K\left(R_0 - R_h\right). &(15)\end{aligned}$$

Constraint (14) says the manager has no incentive to lie when the true state is 0. Constraint (15) says the manager has no incentive to lie when the true state is h. The bank's break-even constraint is

$$D \quad = \quad \pi_h K(R_h - \gamma) + \pi_0(KR_0 - C_0). \tag{16}$$

Condition (12) ensures a monitoring equilibrium is optimal, and the optimal contract satisfies (14) with equality. Solving for the optimal leverage ratio gives

$$D/K \quad = \quad R_h + \rho\pi_0(R_0 - R_h) - \pi_h\gamma. \tag{17}$$

When monitoring is completely ineffective ($\rho = 0$) and free ($\gamma = 0$), (17) reduces to the case without monitoring technology (see equation (11)).

E. Model Predictions

We interpret the staggered implementation of the IBBEA from 1994 to 2005 as a shock to ρ, the banks' probability of learning the true level of profits when monitoring. Expression (17) implies the following testable predictions.

Prediction 1 *Inflexible-price firms have lower leverage than flexible-price firms.*

The expression for leverage (17) increases with firm profits in the low-profit state, R_h. Because inflexible-price firms have lower R_h, leverage decreases with price inflexibility h.

Prediction 2 *Following an increase in the effectiveness of monitoring, inflexible-price firms increase leverage more than flexible-price firms.*

Higher price inflexibility h implies a larger gap between high and low profits, $R_0 - R_h$. Expression (17) then implies leverage increases more in monitoring effectiveness ρ for inflexible-price firms than for flexible-price firms.

III Data

A. Micro Pricing Data

To test the predictions of our model, we use the confidential micro pricing data underlying the Purchaser Price Index (PPI) from the Bureau of Labor Statistics (BLS). We have monthly price information for individual goods at the establishment level from 1982 to 2014. The BLS defines prices as "net revenue accruing to a specified producing establishment from a specified kind of buyer for a specified product shipped under specified transaction terms on a specified day of the month." Unlike the Consumer Price Index

(CPI), the PPI measures the prices from the perspectives of producers. The PPI tracks prices of all goods-producing industries such as mining, manufacturing, and gas and electricity, as well as the service sector.[2]

The BLS follows a three-stage procedure to select their sample of goods. First, it compiles a list of all firms filing with the Unemployment Insurance system to construct the universe of all establishments in the United States. Then, it selects probabilistically sample establishments and goods based on the total value of shipments, or on the number of employees. The final data set covers 25,000 establishments and 100,000 individual items each month. Prices are collected through a survey, which participating establishments receive via email or fax. Individual establishments remain in the sample for an average of seven years, and then a new sample is selected to account for changes in the industry structure.

To compute our measure of frequency of price adjustment (FPA), we first calculate the frequency of price adjustment at the good level as the ratio of price changes to the number of sample months. For example, if an observed price path is $4 for two months and then $5 for another three months, one price change occurs during five months, and the frequency of price adjustment is 1/5. We exclude price changes due to sales.[3] We then equally weigh and aggregate the frequencies to the establishment level via internal identifiers of the BLS. To perform the firm-level aggregation, we check whether establishments with the same or similar names are part of the same company. In addition, for all firms in the data set, we use publicly available data to search for names of subsidiaries and name changes due to, for example, mergers, acquisitions, or restructuring occurring during the sample period.[4]

The granularity of the data at the firm level allows us to differentiate the effect of price flexibility from that of other industry- and firm-level characteristics.

The price flexibility of similar firms operating in the same industry can be substantially different. This difference can arise from different costs of negotiating with customers and suppliers, physical costs of changing prices, or managerial costs such as information gathering, decision making, and communication (see Zbaracki et al.

[2]The BLS started sampling prices for the service sector in 2005. The PPI covers about 75% of the service sector output.

[3]This assumption is standard in the literature and does not change results, because sales are rare in the PPI micro-data. See Gorodnichenko and Weber (2015).

[4]See Weber (2015) for a more detailed description of the data and construction of variables.

(2004)). This paper will not be able to assess the relevance of alternative determinants of price flexibility, which is still a heavily debated topic in Macroeconomics and Industrial Organization.

B. Financial Data

To construct measures of firm-level financials, we focus on firms that have been part of the S&P500 during our sample period from January 1982 to December 2014 due to the availability of the PPI micro data.[5] The S&P500 contains large U.S. firms and captures approximately 80% of the available stock market capitalization in the United States, therefore maintaining the representativeness for the whole economy in economic terms. The BLS samples establishments based on the value of shipments, and we have a larger probability of finding a link between BLS pricing data and financial data when we focus on large firms.

We have 1,195 unique firms in our sample due to changes in the index composition during the sample period, out of which we were able to merge 469 with the BLS pricing data.

Stock returns and shares outstanding come from the monthly stock return file from the Center for Research in Security Prices (CRSP). Financial and balance-sheet variables come from Compustat. $Lt2A$ is defined as long-term debt to total assets; $Prof$ is operating income over total assets; $size$ is the log of total assets; BM is the book-to-market ratio; $It2A$ is intangible assets defined as total assets minus the sum of net property, plant, and equipment, cash and short-term investments, total receivables, and total inventories to total assets; PCM is the price-to-cost margin defined as the ratio of net sales minus the cost of goods sold to net sales. Note this measure is equivalent to 1 minus the operating leverage of the firm; and HHI is the Herfindahl-Hirschman index of sales at the Fama & French 48-industry level.

We follow Lemmon, Roberts, and Zender (2008) and Graham, Leary, and Roberts (2014) in the choice and definition of capital-structure determinants. To reduce the effects of outliers, we winsorize all variables at the 1st and 99th percentiles.

[5]Gorodnichenko and Weber (2015), Weber (2015), and Jaimovich, Rebelo, and Wong (2015) have used similar data.

C. Descriptive Statistics

Panel A of Table 1 reports descriptive statistics for our running sample. Firms in our sample do not adjust their output prices for roughly seven months ($-1/(log(1-FPA))$), with substantial variation across firms as indicated by the large standard deviation. The average long-term leverage ratio $Lt2A$ is around 21%. Firms have an operative income margin ($Prof$) of 15%. Twenty-one percent of assets are intangible ($It2A$). The average book-to-market ratio is 60% (BM), and the average firm size is USD 3.8 bn. ($size$). The average price-to-cost margin (PCM) is 37%, and average industry concentration (HHI) is 0.11. Panel B of Table 1 reports correlations of the variables.

Flexible-price firms have unconditionally higher long-term leverage, and the frequency of price adjustment is unconditionally correlated with standard determinants of capital structure. The frequency of price adjustment is lower in more concentrated industries and for firms with high markups and might, therefore, reflect more market power on the side of firms. In our multivariate analysis, we will keep constant firm-level measures of market power.

D. Are Inflexible-price Firms Riskier?

When firms cannot adjust prices to changing market conditions, cash-flow volatility and profit volatility increase, increasing default risk for a given leverage ratio. To assess the relation between price stickiness and default rates empirically, we obtain default and credit-rating information from Moody's Default and Recovery Database (DRD) and match it to firms in our sample. We construct five default-indicator variables $Default_{t+s}$ for s running from 1 to 5. This dummy is equal to 1 if at least one default occurs within the next $t+s$ years, and 0 otherwise.

Table A.1 in the Online Appendix proposes logistic regressions of default probabilities on the frequency of price adjustment, controlling for firm leverage. Higher leverage is associated with higher default rates. Controlling for leverage, we see that firms with more flexible output prices are less likely to default. The relation between FPA and two- to five-year default rates is statistically significant. The stronger predictive power for multi-year default probabilities is consistent with our observation that uncertainty about the aggregate price level, and hence price rigidities, are likely more relevant over multi-year time horizons.

E. Financial Dependence and Bank Debt

Our sample includes firms in the S&P500 from January 1982 to December 2014, for which we can observe the micro-pricing data. Because our model and our empirical application exploit a shock to bank-level debt, we need to verify that the firms in our sample depend on bank debt rather than only public bond markets. Colla et al. (2013) report that bank loans and credit lines jointly account for at least 30% of the leverage for the largest Compustat firms. This fact suggests that bank debt is an important source of financing for firms with similar characteristics to the ones in our sample.

To assess whether the firms in our sample depend on bank debt, we use the data on credit lines collected by Sufi (2009).[6] These data allow observation of an extensive margin of credit lines—whether firms have an active credit line or not—and an intensive margin of credit lines—the share of the line that has been used at each point in time. We can construct the extensive margin for all the firm-year observations in our sample, whereas the intensive margin is only available for those firms that match with the 5% random sample of Compustat firms constructed by Sufi (2009).

As for the extensive margin, the vast majority of the firm-year observations in our sample have a credit line open with at least one bank (94.6%). Consistent with our model's prediction, we find flexible-price firms are more likely to have a credit line (97.3%) than inflexible-price firms (93.6%), and a t-test for whether these ratios are equal rejects the null at the 1% level of significance. Moving on to the intensive margin, we find the usage rate of credit lines for firms in our sample is 24.8%. An economically significant difference exists in the usage rate across inflexible-price firms (28.1%) and flexible-price firms (15.6%). A t-test for whether these ratios are equal rejects the null at the 5% level of significance. In Figure A.1 of the Online Appendix, we plot the density of the usage ratio for the two groups of firms. The full distribution of the usage ratio for inflexible-price firms lies to the right of the distribution for flexible-price firms. Although inflexible-price firms are less likely to have a credit line with banks, they are more likely to draw down the credit line, indicating they might be more credit constrained than flexible-price firms.

[6] In contrast to Capital IQ, Sufi (2009) has comprehensive coverage starting in 1996 and information on drawn and undrawn credit lines.

IV Baseline Analysis

A. Price Flexibility and Leverage

We move on to investigate the empirical relationship between leverage and price stickiness. Inflation is highly persistent (Atkeson and Ohanian (2001), Stock and Watson (2007)), and uncertainty about the aggregate price level increases with the forecast horizon. Price-setting frictions should therefore be most relevant for profits over long horizons. In addition, Heider and Ljungqvist (2015) argue firms use short-term leverage to finance working capital, and are therefore unlikely to change short-term leverage in response to changing tax benefits or credit supply. For these reasons, we focus on long-term debt, as opposed to short-term debt, in our empirical analysis. In the Online Appendix, we replicate all the results using total debt as the outcome of interest. The results are qualitatively similar, although the sizes of the associations are lower, consistent with the prediction that price-setting frictions are more relevant for profits over long horizons.[7]

As a first step, we plot the correlation between long-term debt to assets and the price flexibility of firms over time. In both panels of Figure 1, the blue solid lines refer to the ratio of long-term debt to assets of firms in the bottom quartile by price flexibility. The red long-dash lines refer to the ratio of long-term debt over assets of firms in the top quartile by price flexibility, and the black short-dash lines are the difference between the two ratios. In both panels, the baseline prediction of our model seems satisfied: flexible-price firms have on average higher long-term leverage than inflexible-price firms, and this fact is true over the whole time period our sample covers.

In the top panel of Figure 1, the red vertical line indicates 1996, which is the year the first set of U.S. states started to implement the Interstate Bank Branching Efficiency Act (IBBEA), an event we describe and exploit for our identification strategy below. In the bottom panel of Figure 1, the red vertical line indicates 2000, which is the year the last group of U.S. states that had not started to implement the IBBEA between 1996 and 1998 indeed started the implementation. In both panels, the difference in the ratio of long-term debt to assets is stable to the left of the vertical lines, and it declines to the right of the vertical lines. We will exploit these events and the convergence of the ratios for the two groups of firms below to test for the second prediction of our model.

[7]See Table A.4, Table A.5, and Table A.6 in the Online Appendix for the results using total debt.

B. Ordinary Least-Squares Analysis

To assess the magnitude of the correlation between price flexibility and long-term debt to assets, our most general specification is the following OLS equation:

$$Lt2A_{i,t} = \alpha + \beta \times FPA_i + X'_{i,t-1} \times \gamma + \eta_t + \eta_k + \epsilon_{i,t}, \tag{18}$$

where $Lt2A_{i,t}$ is long-term debt to assets of firm i in year t; FPA is the frequency of price adjustment, which is higher for firms with more flexible prices; X is a set of standard determinants of capital structure; η_t is a set of year fixed effects, which absorb time-varying shocks all firms face, such as changes in economy-wide interest rates; η_k is a set of industry fixed effects, which absorb time-invariant unobservable characteristics that differ across industries. We use two definitions of industry fixed effects. The coarser definition allows for variation within 1-digit SEC codes, which splits the firms in our sample into eight industry groups. The tighter definition allows for variation within the 48 Fama-French industries.[8] Across all specifications, we cluster the standard errors at the firm level to allow for correlation of unknown form across the residuals of each firm over time.

Table 2 displays the baseline results. In columns (1)-(3), FPA is the continuous measure of price flexibility, whereas in columns (4)-(6), it is a dummy variable that equals 1 for the firms in the top 25% of the distribution based on price flexibility, and 0 for the firms in the bottom 25% of the distribution.

In column (1) of Table 2, we regress the ratio of long-term debt to assets on price flexibility and standard determinants of capital structure. Firms with more flexible output prices have a higher ratio of long-term debt over total assets. This positive association is significantly different from 0 at the 1% level of significance. A one-standard-deviation increase in price flexibility (0.14) is associated with 2.1-percentage-point increase in the ratio of long-term debt to assets, which is 10% of the average ratio in the sample. In columns (2)-(3), we only exploit variation in the frequency of price adjustment across firms within the same year, and across firms within the same industry, and we confirm the results in column (1).

[8]Note we cannot restrict the variation within firm, because the measure of frequency of price adjustment is time invariant. As we show below, even when we measure the frequency of price adjustment in different subsamples of the data, the correlation of the variables at the firm level is statistically indifferent from 1.

In columns (3)-(6), we estimate the analogous specifications using the indicator for firms with the most flexible prices, and look only at the most flexible firms (top 25% of the distribution by price flexibility) and the least flexible firms (bottom 25% of the distribution by price flexibility). We confirm the results we obtained with the continuous measure of price flexibility.[9] Being in the top quarter of the distribution of firms by price flexibility is associated with a 6-percentage-point higher ratio of long-term debt over assets. The results are qualitatively similar when we only exploit within-year and within-industry variation in price flexibility across firms.

In untabulated results, we find the correlation between price flexibility and leverage does not change if we add other firm-level controls in equation (18), for instance, when we control for cash over assets, following Faulkender et al. (2006), who show cash flows are an important determinant of firm-level leverage targets and the speed of adjustment toward such targets.

C. Errors-in-Variables Specifications

Erickson, Jiang, and Whited (2014) propose a novel methodology to account for the measurement error in explanatory variables using linear cumulant equations. They show several firm-level determinants of capital structure change sign or lose statistical significance once they correct for measurement errors. We follow their methodology to assess the robustness of the association between price flexibility and long-term leverage when correcting for measurement error in key variables. Specifically, we follow Erickson et al. (2014) assuming measurement error possibly affects two key determinants of capital structure: asset intangibility and the book-to-market ratio. In addition, we also assume the measure of price flexibility is measured with error. This assumption seems plausible, because the measure is based on the aggregation of frequencies of price adjustment at the good level based on a representative sample of goods. Note the errors-in-variables estimator we compute will therefore assume that the other variables in our specification are not measured with error.

In column (1) of Table 3, we report the baseline OLS estimator from column (1) of Table 2 to allow comparison across estimators. In columns (2)-(4), we report the estimated

[9]The results are similar when we add all other firms and assign them a value of 0 for the FPA dummy measure.

coefficients when implementing the cumulant-equation method of Erickson et al. (2014) for the third, fourth, and fifth cumulants. We do not report the results for using higher-order cumulants in the linear equations, because of the size of our sample. Using higher-order cumulants results in estimates of similar size and substantially lower standard errors. Comparing the estimated association of price flexibility with long-term leverage across specifications, the size and significance of the coefficients are similar in the baseline OLS specification and when using the errors-in-variables estimator. The results for the other covariates are in general similar, but some lose statistical significance or switch sign, including the two covariates we also assume are measured with error (book-to-market ratio and asset intangibility).

V Identification Strategy and Falsification Tests

To assess whether the effect of price flexibility on leverage is causal, one route would be testing the effect of a shock to firm-level price flexibility on leverage, or proposing an instrument for price flexibility. Price flexibility is a highly persistent characteristic of firms. For instance, in our sample, a firm-level regression of post-1996 price flexibility onto pre-1996 price flexibility yields a slope coefficient of 93%, and we fail to reject the null that the coefficient equals 1 at any plausible level of significance.[10] This persistence suggests we can hardly consider a shock to firm-level price flexibility for identification purposes in our sample.

We therefore propose an identification strategy derived from our model's predictions, and in line with the financial-constraints literature. We (i) identify a positive shock to the quality of monitoring of the banks firms can access, (ii) show inflexible-price firms increase leverage more than flexible-price firms, and (iii) show the effect does not revert in the short run. Our strategy exploits a quasi-exogenous shock to financial constraints, and uses ex-ante unconstrained firms to assess the causal effect of financial constraints on inflexible-price firms.

To implement this strategy, we need a quasi-exogenous shock to firm-level financial constraints, as well as a viable control group of firms to assess how inflexible firms' long-term leverage would have evolved absent the shock.

[10] See also Nakamura and Steinsson (2008), Golosov and Lucas (2007), and Alvarez et al. (2011).

The shock we use is the staggered state-level implementation of the Interstate Banking and Branching Efficiency Act (IBBEA) of 1994. The IBBEA represented a shock to the ability of banks to open branches and extend credit across state borders. This shock is relevant for the leverage of firms in our sample, because in section III.E, we find 95% of them have a credit line open with at least one bank, and all firms use such lines, especially the inflexible-price firms (see Figure A.1 in the Online Appendix).

For the control group, we use flexible-price firms in the same states and the same years as inflexible-price firms to proxy for the behavior of inflexible-price firms absent the shock. To assess the plausibility of the parallel-trends assumption we make, below we show the pre-shock trends of long-term leverage for inflexible- and flexible-price firms are similar, and the change in the price flexibility of firms in our sample before and after the shock is economically and statistically indifferent from zero.

A. Institutional Details and Interpretation

Restrictions to banks' geographic expansion have a long history in the United States (Kroszner and Strahan (2007)). The McFadden Act of 1927 gave states the authority to regulate in-state branching, and most states enforced restrictions on branching well into the 1970s. In 1970, only 12 states allowed unrestricted in-state opening of branches, and 16 states prohibited banks from opening more than a single branch. In addition to branching restrictions, the Douglas Amendment to the 1956 Bank Holding Company Act effectively prohibited a bank holding company from acquiring banks outside the state where it was headquartered (Strahan (2003)).

Starting in the 1970s, the restrictions on acquiring banks across states were gradually eased. Kroszner and Strahan (1999) argue the timing of this deregulation wave relates to the interaction of technological innovations, such as the ATM, with lobbying by large and well-capitalized banks, but not to time-varying local economic conditions. Instead, before the Interstate Banking and Branching Efficiency Act (IBBEA) of 1994 banks needed the target state's explicit approval to open branches across state lines.

The approval of IBBEA was a watershed event for interstate banking, but did not immediately lead to nationwide branching in all states. The law permitted states to (a) require a minimum age of the acquired institution, (b) restrict *de novo* interstate branching, (c) disallow the acquisition of individual branches without acquiring the entire bank, and (d) impose statewide deposit caps. We use Rice and Strahan (2010)'s time-varying index for regulatory constraints between 1994 and 2005 to construct a dummy variable that equals 1 in the year the state lifted at least one of the restrictions (a) through (d), and in all the subsequent years. In the following sections, a state is deregulated when this dummy variable equals 1, and it is not deregulated otherwise.[11]

B. Triple-Differences Strategy

We propose a triple-differences strategy. Our strategy exploits the time variation in the exposure to the implementation of the IBBEA firms faced based on their location. Moreover, we use the group of most flexible-price firms as a counterfactual for the evolution of long-term debt of inflexible-price firms absent the deregulation shock. The idea is that flexible-price firms were not borrowing constrained before 1996, because they were less risky compared to inflexible-price firms.

B.1 Parallel-Trends Assumption

A necessary condition for identification is the *parallel-trends assumption*, which states that the evolution of long-term debt of flexible- and inflexible-price firms would have followed common trends before *and* after the shock, had the shock not happened. The potential outcome absent the shock is unobservable, and hence we cannot test this assumption directly. At the same time, we can assess the extent to which the trends of long-term leverage across flexible- and inflexible-price firms are parallel before the shock. If we are convinced the pre-trends are parallel, our identifying assumption would be that any divergence in the trends after the shock is due to the shock itself, and not to other possible concurrent shocks or alternative explanations. Under this identifying assumption, the evolution of long-term debt of flexible-price firms represents a valid counterfactual to the evolution of long-term debt of inflexible-price firms had they not been exposed to the deregulation.

Figure 2 proposes a graphical assessment for whether the trends in long-term leverage are parallel across flexible- and inflexible-price firms in the years before the first states implement the IBBEA, which is 1996. Figure 2 plots the estimated coefficients $\hat{\beta}_t$ and the

[11]No states re-instated any restriction they had already lifted. Several states lifted the restrictions (a) through (d) in different years from 1996 until 2002.

95% confidence intervals from the following OLS specification:

$$
\begin{aligned}
Lt2A_{i,t} = \alpha + \sum_{t=1983}^{1996} \beta_t \times FPA_i \times \eta_t \\
+ \delta_1 \times FPA_i + \eta_t + \epsilon_{i,t},
\end{aligned}
\tag{19}
$$

which includes a set of leads of the interactions between price flexibility and year fixed effects for the years before the first IBBEA implementations (1996). The excluded year is 1982. The estimated coefficient $\hat{\delta}_1$ equals 0.092 (t-stat 5.54), and statistical inference is based on standard errors clustered at the firm level. The size of the confidence intervals is similar if we allow for correlation of unknown form across observations in a same state, as opposed to a same firm. We fail to reject the null hypothesis that any of the interaction terms between price flexibility and year fixed effects is different from zero in all years before the first implementations of the IBBEA except 1995, when the interaction is positive.

B.2 Price Flexibility around the Shock

A large body of research in Macroeconomics finds the extent of price flexibility is a highly persistent feature of firms (e.g., see Golosov and Lucas (2007) and Alvarez et al. (2011)). Ideally, we would like to test formally that the price flexibility of the firms in our sample did not change over time, and especially that the implementation of the IBBEA at the state level did not affect it. The construction of the measure of price flexibility based on the BLS micro-pricing data does not allow computing yearly values, because we need to observe several price spells for a given good for the measure to be meaningful.

Therefore, we proceed as follows. We identify the firms in our sample for which we can observe monthly price spells for the three years before and after 1996. We construct a measure of price flexibility before 1996, based on the monthly spells in the period 1993-1995, and a measure of price flexibility after 1996, based on the monthly spells in the period 1996-1998. We then regress the post-1996 measure on the pre-1996 measure and a constant. Our null hypothesis is the regression coefficient equals 1; that is, the pre-1996 measure is perfectly correlated with the post-1996 measure. Our estimated coefficient equals 0.93, and we cannot reject the null that this coefficient differs from 1 at any plausible level of significance. The 95% confidence interval around the coefficient is (0.73; 1.12). Also note we truncate price spells by only focusing on a three-year period, and hence we introduce noise in our measures. The almost perfect correlation in the frequency of price adjustment before and after 1996 is therefore hardly consistent with the notion that firm-level price flexibility changed around the implementation of the IBBEA.

B.3 Triple-Differences Specification

To implement our strategy, we estimate the following specification:

$$\begin{aligned} Lt2A_{i,t} = \alpha + \beta \times FPA_i \times Deregulated_{i,t} \\ + \delta_1 \times FPA_i + \delta_2 \times Deregulated_{i,t} + \eta_t + \eta_f + \epsilon_{i,t}, \end{aligned} \tag{20}$$

where $Deregulated_{i,t}$ is an indicator that equals 1 if firm i is in a state that had implemented the deregulation in year t, and 0 otherwise; and η_t and η_f are a full set of year and industry fixed effects.

Equation (20) compares the long-term debt raised within firms before and after their state implemented the deregulation, across firms in deregulated and regulated states, and across flexible- and inflexible-price firms.

Based on our model, we have the following predictions on the coefficients in equation (20): $\delta_1 > 0$, because, on average, higher price flexibility leads to more long-term debt; and $\delta_2 \geq 0$, because firms have more funds available to borrow after the 1994 deregulation shock, which could be 0 because flexible-price firms were unlikely to be financially constrained before the shock. The crucial prediction of our strategy is that $\beta < 0$, because the most inflexible-price firms obtain disproportionally more funds after the deregulation compared to the most flexible-price firms.

Table 4 reports the estimates for the coefficients in equation (20). In columns (1)-(3), FPA is the continuous measure of price flexibility; in columns (4)-(6), it is the dummy that equals 1 for firms in the top 25% of the distribution based on price flexibility, and 0 for those in the bottom 25% of the distribution. For both sets of results, the first column reports estimates for the baseline specification. In the second column, we add year fixed effects and eight industry-level dummies, which capture the one-digit SIC industry to which the firm belongs. In the third column, we add year fixed effects and the 48 industry-level dummies for the Fama-French industry taxonomy.

Across all specifications, the sign of the estimated coefficients are in line with the predictions above. Firms with higher price flexibility have higher long-term debt on average ($\hat{\delta}_1 > 0$). More importantly, across all specifications, we find that flexible-price firms increase their leverage less than inflexible-price firms after the state-level implementation of the deregulation ($\hat{\beta} < 0$). The effect of the deregulation on the most flexible-price firms is close to zero, as can be seen by adding the estimates of $\hat{\beta}$ and $\hat{\delta}_1$ across all the specifications. Comparing column (1) with columns (2)-(3), and column (4) with columns (5)-(6), shows that the size of the estimated interaction effect does not change when we only exploit within-industry variation. Therefore, whereas industry-level effects explain about half of the size of the baseline effect of price flexibility on leverage, the variation across firms within the same industries explains the full size of the effect of financial constraints across flexible- and inflexible-price firms.

For the purposes of statistical inference, in Table 4, we cluster the standard errors at the level of the firm. All the t-statistics are higher - that is, the estimated standard errors are lower - if we instead cluster the standard errors at the level of the state, which is the level at which our treatment is administered. The most plausible explanation for this result is the specifications that allow for correlation of unknown form within the same states have few clusters, because we only observe firms in 42 different states. For this reason, and because we cluster standard errors at the firm level in our baseline analysis, we report the results corresponding to the more conservative standard errors clustered at the firm level.

B.4 Effect at Impact and Over Time

Our tests so far have used observations for a same firm in different years, both before and after the implementation of the IBBEA. Bertrand et al. (2004) show that the autocorrelation between observations of a same unit over time might understate dramatically the size of the standard errors in difference-in-differences research designs. We tackle this issue in Table 5. First, we estimate equation (20) using only two data points for each firm. We only keep firm-level observations in the year before the deregulation and the year after the deregulation is implemented in their state. For each firm, we only have two observations, one of which is before the shock and one after the shock. This test aims to estimate the effect of the shock at impact, that is, around the year in which the shock happened. We report the results for this test in column (1) of Table 5. We have 599 observations; hence, the size of the sample is drastically lower than our baseline sample in Table 4. The results are qualitatively similar to those in Table 4, although the size of the estimated coefficient is about half the size of the corresponding coefficient in column (1) of Table 4.

In columns (2)-(5) of Table 4, we report the results for estimating equation (20) in periods of different length. In column (2), we only use observations from 1994 until 2002, which includes the years in which the first and the last state implemented the IBBEA (1996 and 2001, respectively). In each of columns (3)-(5), we enlarge the time period by three years going backward and forward. Qualitatively, our results are similar across these different time periods. Interestingly, the size of the interaction between price flexibility and the IBBEA implementation increases monotonically in absolute value as long as we add observations in later years. At the same time, the baseline effect of price flexibility on leverage stays the same across the sub-periods. These results are consistent with the notion that the IBBEA implementation could not be immediate on either the side of the banks or of the firms. The process of opening new branches and consolidating must have developed over time after a state implemented the IBBEA. At the same time, firms take time to adjust their capital structure, that is, to change the contracts that regulate the sources of financing of the firm. Note diverging trends between flexible- and inflexible-price firms before the shock cannot be driving these results, because we showed the trends are parallel before the shock in Figure 2.

B.5 Effect by dependence on external financing

To corroborate the interpretation of the deregulation shock, we exploit the cross-sectional variation in terms of the financial dependence of firms. If the banking-deregulation shock is truly driving the interaction effect, then inflexible-price firms that depend more on external financing to fund their operations should drive this effect. We thus estimate the specification in equation (20) separately for firms in the top tercile by cash-to-assets ratio and for other firms. The rationale is that inflexible-price firms with high cash-to-assets ratios will not depend much on external financing, whereas inflexible-price firms with lower cash-to-assets ratios should be those the deregulation shock affects. Consistent with this interpretation, Table 6 shows the effect of deregulation on firms' leverage is driven by inflexible-price firms with low cash-to-assets ratios (columns (1) and (3)), as opposed to those with high cash-to-assets ratios (columns (2) and (4)).

B.6 Robustness

In the Online Appendix, we show the results of our triple-differences specifications are robust across subsamples and using alternative definitions of price flexibility. Table A.2 in the Online Appendix replicates all the results when we exclude financial firms and utilities. Table A.3 estimates our main specification from equation (20) separately within one-digit SIC industries. The results are similar across all industries except utilities and finance.

C. Falsification Tests

To further assess the validity and interpretation of our causal test, we propose an empirical setup that allows the design of two falsification tests (Roberts and Whited, 2013). We exploit the fact that the state-level implementation of the IBBEA was not only staggered over time, but also clustered in two periods. The majority of U.S. states implemented the deregulation between 1996 and 1998. The second group of states only implemented the deregulation after 2000. We call the first group of states "early states," and the second group, "late states." This setup allows us to construct three tests across three groups of years. Before 1996, no state had implemented the deregulation yet. Between 1996 and 2000, firms in early states were exposed to the deregulation, but firms in late states were not. After 2000, all firms were in states that had deregulated. Figure 3 gives a graphical depiction of this setup.

In a first specification, we corroborate our triple-differences result in the novel setup, by estimating the following:

$$\begin{aligned} Lt2A_{i,t} = \alpha &+ \beta \times FPA_i \times After1996_{i,t} \times EarlyState_i + \delta_1 \times FPA_i \times After1996_{i,t} \\ &+ \delta_2 \times FPA_i \times EarlyState_i + \delta_3 \times After1996_{i,t} \times EarlyState_i + \gamma_1 \times FPA_i \\ &+ \gamma_2 \times After1996_{i,t} + \gamma_3 \times EarlyState_i + X'_{i,t} \times \zeta + \epsilon_{i,t}. \end{aligned} \tag{21}$$

Panel A of Figure 3 sketches the predictions of our model for the specification in equation (21). It is a quadruple-differences design, because it compares outcomes within firms before and after 1996, across firms before and after 1996, across firms in early and late states, and across flexible- and inflexible-price firms. To corroborate our earlier results, we estimate equation (21) using only firm-level observations up to 2000. The rationale is that firms in early states were exposed to the deregulation between 1996 and 2000, whereas firms in late states were not. Flexible- and inflexible-price firms in late states thus represent the control group for the differential evolution of long-term debt in flexible- and inflexible-price firms in early states, had they not been exposed to the deregulation shock.

Our prediction is that $\beta < 0$, $\delta_1 = 0$, and $\gamma_1 > 0$; that is, flexible-price firms have higher leverage on average, and after the deregulation, only inflexible-price firms in early states increase their leverage compared to flexible-price firms in early states. The baseline effect of price flexibility on leverage should not change after 1996 for firms in late states.

The estimates in column (1) of Table 7 support our predictions. In columns (2)-(3) of Table 7, we repeat the analysis separately for firms with low and high cash-to-assets ratios. Similar to our earlier results, the subsample of firms with a higher need for external financing drive the effects.

We then proceed to assess the validity of our designs by constructing two falsification tests. Panel B of Figure 3 sketches the predictions of our model for our first falsification test. We build on the specification in equation (21), but we limit our estimation to the firm-level observations before 1996. This limitation implies no firm-level observations, either in early or late states, are exposed to the deregulation shock. Because in the baseline analysis we use a treatment period of four years for early states, from 1996 to 2000, we assign 1992 as a placebo deregulation year to observations in early states. We thus replace the dummy $After1996_{i,t}$ in equation (21) with the dummy $After1992_{i,t}$, which equals 1 for all firm-level observations after 1992. Our falsification test consists of comparing flexible- and inflexible-price firms in early and late states after 1992, and before the deregulation happened. If our earlier test was invalid, and our baseline results captured the effect of state-level characteristics different across early and late states, but unrelated to the deregulation event, we should reject the null hypothesis that $\beta = 0$. Column (4) of Table 7 shows that, instead, we fail to reject this null hypothesis at a plausible levels of significance. As expected, we find flexible-price firms have higher leverage on average, irrespective of the states where they are located.

Panel C of Figure 3 sketches the predictions of our model for our second falsification test, in which we exclude all firm-level observations between 1996 and 2000. This limitation implies that in each year, the observations in early and late years are either not exposed to the deregulation shock (before 1996), or they are all exposed to the deregulation shock (after 2000). We thus estimate the same specification in equation (21), but the new setup implies different predictions from those discussed above. On the one hand, we should not be able to reject the null that $\beta = 0$, because early and late states are exposed to the deregulation in the same years. On the other hand, we now do expect $\delta_1 < 0$ and $\gamma_1 > 0$, because flexible-price firms in both early and late states should have on average higher leverage, and should react less than inflexible-price firms to the deregulation shock. We find evidence consistent with these predictions in column (5) of Table 7.

VI Conclusion

We show that firms with inflexible output prices have lower leverage relative to firms with flexible prices, after controlling for standard determinants of capital-structure choices and exploiting variation within industries and within years. We interpret this fact in a costly-state-verification model, in which inflexible-price firms cannot adjust their prices to macroeconomic shocks, and banks can access a costly monitoring technology. The model also predicts inflexible-price firms should increase their leverage by more than flexible-price firms after a positive shock to the monitoring technology of banks. Using the staggered state-level implementation of the 1994 Interstate Bank Branching Efficiency Act, we test for this prediction in a triple-differences strategy.

These results suggest price flexibility is an important determinant of firms' capital structure. Because firm-level price flexibility is highly persistent over time, these results also suggest price flexibility might contribute to an understanding of persistent differences in financial leverage across firms documented for the first time by Lemmon, Roberts, and Zender (2008).

Price rigidity has a long tradition in research across fields as different as Marketing, Industrial Organization, and Macroeconomics. Our results open up exciting avenues for future research at the intersection of Corporate Finance, Macroeconomics, and Industrial Organization.

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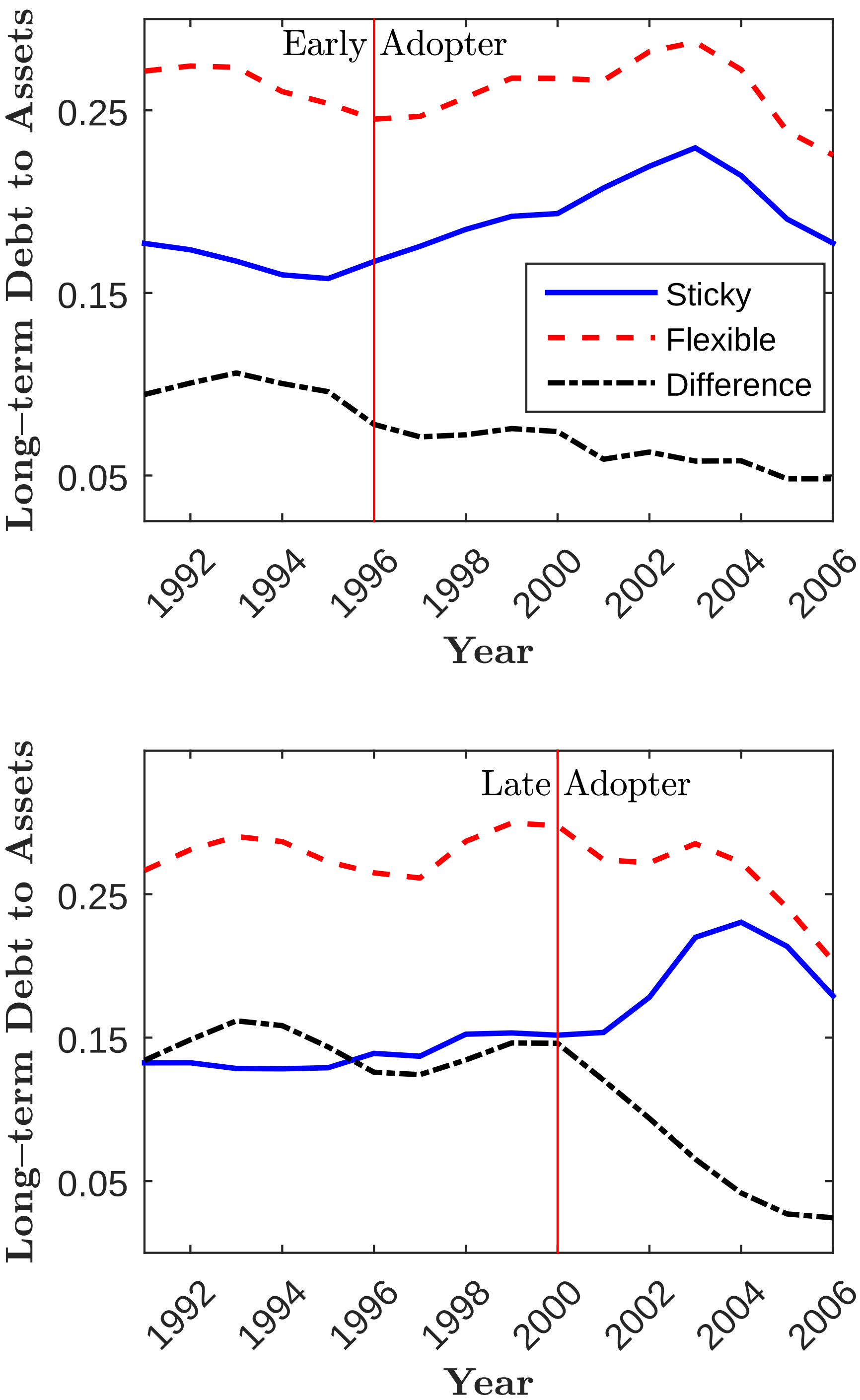
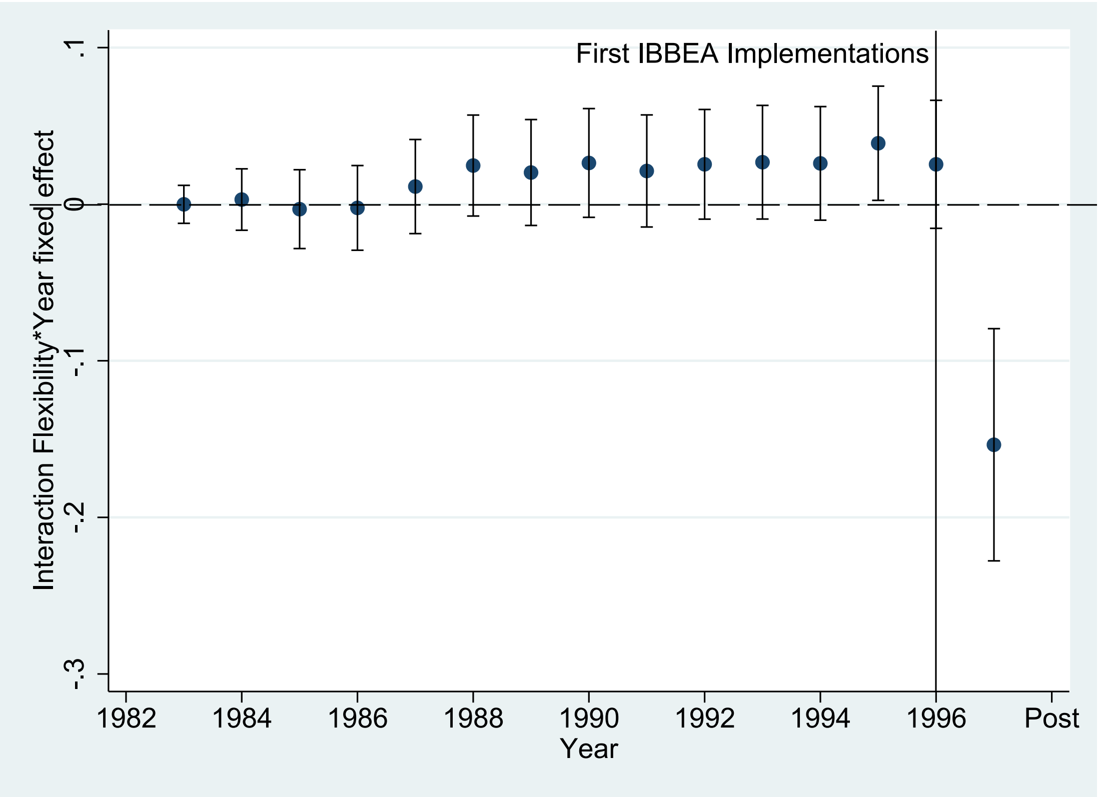
Figure 1: **Long-Term Debt and Price Flexibility**



This figure plots long-term debt to total assets for different percentiles of the price stickiness distribution. Sticky firms are firms below the bottom 25th percentile of the distribution. Flexible-price firms are firms above the top 25th percentile of the distribution. Equally-weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. The sample period is January 1980 to December 2013.

Figure 2: **Parallel Trends Assumption: Assessment of Pre-Trends**

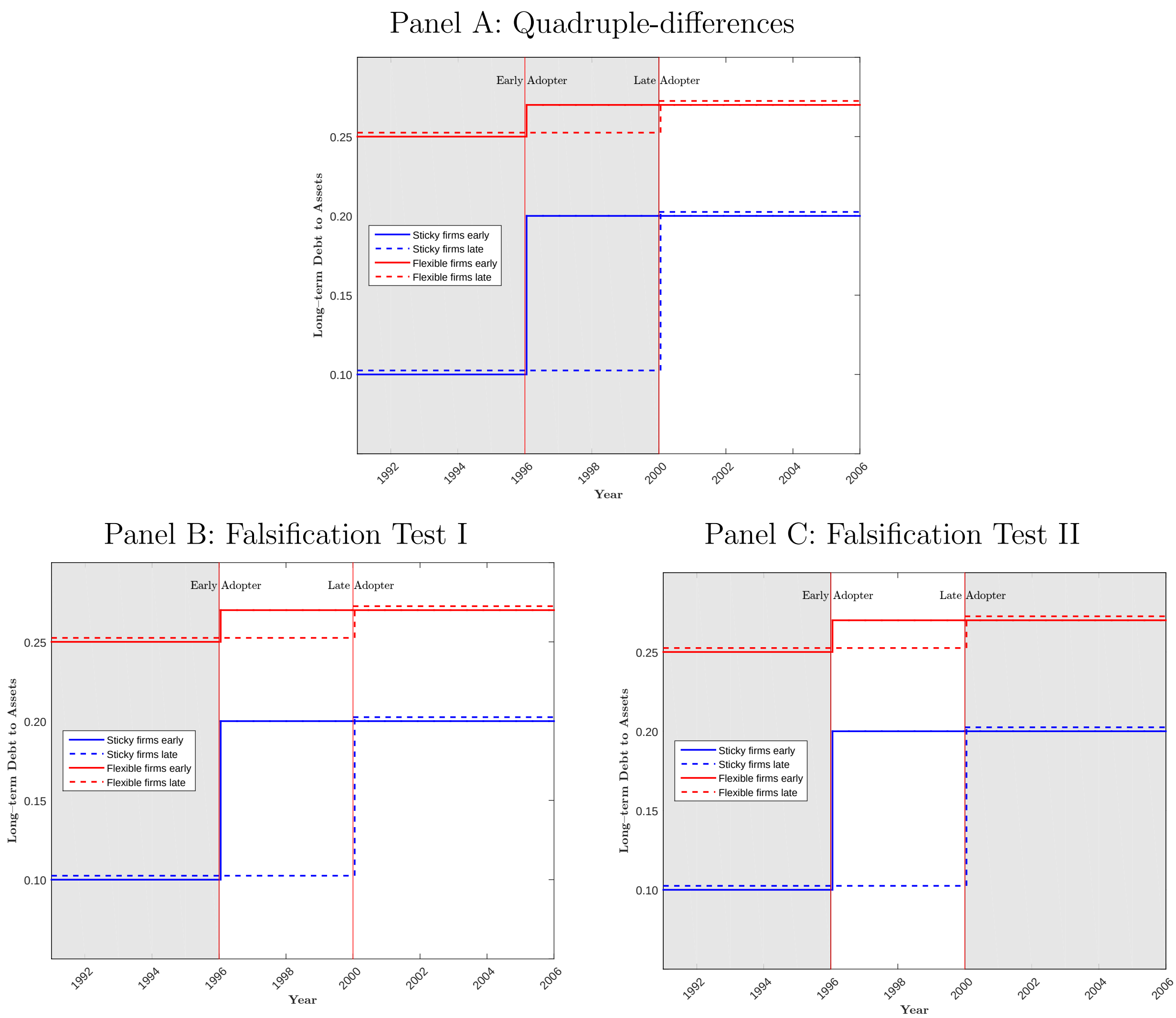


This figure plots the estimated coefficients $\hat{\beta}_t$ and the 95% confidence intervals from the following OLS specification:

$$Lt2A_{i,t} = \alpha + \sum_{t=1983}^{1996} \beta_t \times FPA_i \times \eta_t + \delta_1 \times FPA_i + \eta_t + \epsilon_{i,t},$$

which includes a set of leads of the interactions between price flexibility and year fixed effects for the years before the first IBBEA implementations (1996). The excluded year is 1982. The estimated coefficient $\hat{\delta}_1$ equals 0.092 (t-stat 5.54). Statistical inference is based on standard errors clustered at the firm level.

Figure 3: **Falsification Tests**



This figure describes our quadruple-differences strategy (Panel A) and two falsification tests (Panels B and C). The shaded areas represent the years whose observations we exploit in each test. All Panels report the model's predictions for the evolution of the ratio of long-term debt to assets across four groups of firms in the period 1988 to 2006. In each Panel, the two bottom lines refer to inflexible-price firms in early states that implemented the deregulation of interstate branching between 1996 and 1998 (blue, solid), and in late states that implemented the deregulation after 2000 (red, dashed). The two top lines refer to flexible-price firms in early states (black, dotted) and late states (green, dash-dot). The model predicts in each type of state, the increase in the ratio of long-term debt to assets increases more for inflexible-price firms than for flexible-price firms after the deregulation. In Panel A, we only use observations up to 2000. We therefore propose a quadruple-differences strategy, whereby we compare the outcome within firms before and after 1996, across firms before and after 1996, between early and late states, and between flexible- and inflexible-price firms. The model predicts firms in early states increase their long-term debt to assets in 1996, whereas firms in late states do not. Moreover, inflexible-price firms in early states increase their long-term debt to assets more than flexible-price firms in 1996. In Panel B, we depict the first falsification test, where we only use observations up to 1996. Before 1996, no firms where exposed to the deregulation, and hence the model predicts no differences in long-term to assets across firms in early and late states. Instead, the model predicts the baseline difference in leverage between flexible- and inflexible-price firms, irrespective of their state. In Panel C, we depict the second falsification test, where we use observations before 1996 and after 2000, and hence we exclude the period 1996-2000. In this case, either all firms are in deregulated states, or they are in regulated states. Thus, the model predicts no difference in the change in leverage across firms in early and late states. It predicts the baseline difference in leverage across flexible- and inflexible-price firms, as well as the larger increase in leverage for flexible-price firms at the deregulation event.

Table 1: **Summary Statistics**

This table reports descriptive statistics for the variables used in the empirical analysis for firms with non-missing price stickiness measure in Panel A, and correlations across variables in Panel B. FPA measures the frequency of price adjustment. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. $FPADummy$ is a dummy that equals 1 for firms in the top 25% of the distribution by the frequency of price adjustment; $Lt2A$ is long-term debt to total assets; $Prof$ is the profitability ratio, defined as operating income before depreciation to sales; $Cf2A$ is income and extraordinary items plus depreciation and amortization to total assets; $C2A$ is cash and short-term investments to total assets; $It2A$ is intangible assets to total assets; PCM is the price-to-cost margin; and HHI is the Herfindahl-Hirschman index of sales at the Fama & French 48 industry level. Stock-level data are from CRSP and financial statement data are from Compustat. The sample period is January 1982 to December 2014.

Panel A. Summary Statistics

	FPA (1)	FPA Dummy (2)	Lt2A (3)	Prof (4)	Size (5)	BM (6)	It2A (7)	PCM (8)	HHI (9)
Mean	0.14	0.25	0.21	0.15	8.25	0.60	0.26	0.37	0.11
Median	0.07	0.00	0.20	0.15	8.29	0.49	0.23	0.34	0.08
Std	0.14	0.43	0.13	0.08	1.39	0.41	0.17	0.18	0.10
Min	0.00	0.00	0.00	−0.47	4.16	0.05	0.01	0.05	0.01
Max	0.71	1.00	0.62	0.97	11.62	2.23	0.74	0.83	0.93
Nobs	9,133	9,133	9,176	9,182	9,190	9,054	9,105	9,190	9,190

Panel B. Correlations

	FPA (1)	FPA Dummy (2)	Lt2A (3)	Prof (4)	Size (5)	BM (6)	It2A (7)	PCM (8)
FPA Dummy	0.869***							
Lt2A	0.249***	0.200***						
Prof	−0.144***	−0.115***	−0.284***					
Size	0.128***	0.116***	0.121***	−0.0649***				
BM	0.341***	0.274***	0.258***	−0.456***	−0.0253*			
It2A	−0.224***	−0.187***	0.109***	−0.131***	0.297***	−0.192***		
PCM	−0.212***	−0.194***	−0.167***	0.461***	−0.190***	−0.383***	0.141***	
HHI	−0.0925***	−0.0976***	−0.0647***	0.133***	0.133***	−0.164***	0.136***	0.0581***

-

Table 2: **Panel Regressions of Leverage on Price Flexibility**

This table reports the results for estimating the following OLS equation:

$$Lt2A_{i,t} = \alpha + \beta \times FPA_i + X'_{i,t-1} \times \gamma + \eta_t + \eta_k + \epsilon_{i,t},$$

where $Lt2A$ is long-term debt to total assets, FPA is the frequency of price adjustment, $Prof$ is operating income over total assets, $Size$ is the logarithm of sales, BM is the book-to-market ratio, $It2A$ is intangible assets to total assets, PCM is the price-to-cost margin, and HHI is the Herfindahl-Hirschman index of sales at the Fama & French 48 industry level. $IndustryFE1$ is a set of eight dummies that capture the 1-digit SIC codes of firms. $IndustryFE2$ is a set of forty-eight dummies that capture the Fama & French 48 industries. The sample period is January 1982 to December 2014. Stock-level data are from CRSP and financial statement data are from Compustat. Standard errors are clustered at the firm level. Columns (1) to (3) use the continuous measure of the frequency of price adjustment and columns (4) to (6) use a dummy which equals 1 if the firm is in the top tercile of the frequency of price adjustment distribution. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics.

	FPA			FPA Dummy		
	(1)	(2)	(3)	(4)	(5)	(6)
FPA	0.18***	0.07**	0.09***	0.06***	0.02*	0.04**
	(4.96)	(2.33)	(3.10)	(3.93)	(1.75)	(2.41)
Prof	−0.22***	−0.24***	−0.32***	−0.20**	−0.26***	−0.29***
	(−3.02)	(−3.92)	(−5.76)	(−2.39)	(−3.78)	(−4.08)
Size	0.00	0.00	0.00	0.00	0.00	0.00
	(1.33)	(0.40)	(0.33)	(−0.11)	(−0.29)	(−0.07)
BM	0.05***	0.02**	0.02*	0.06***	0.04***	0.03***
	(5.31)	(2.27)	(1.73)	(5.46)	(4.02)	(2.81)
It2A	0.11***	0.12***	0.14***	0.14***	0.15***	0.15***
	(3.83)	(4.21)	(5.07)	(3.23)	(3.66)	(3.84)
PCM	0.00	−0.03	0.06**	0.02	0.00	0.09**
	(−0.12)	(−1.13)	(2.00)	(0.40)	(0.07)	(2.36)
HHI	−0.03	0.07	0.07	−0.03	0.08	0.09
	(−0.72)	(1.65)	(1.42)	(−0.48)	(1.27)	(1.37)
Constant	0.12***	0.22***	0.19***	0.15***	0.22***	0.18***
	(3.47)	(4.90)	(4.59)	(3.63)	(4.23)	(3.66)
Year FE1	X	X	X	X	X	X
Industry FE1		X			X	
Industry FE2			X			X
Nobs	8,824	8,824	8,824	4,408	4,408	4,408
Adjusted R^2	0.16	0.27	0.34	0.18	0.28	0.37

t-stats in parentheses
$*p < 0.10, **p < 0.05, ***p < 0.01$

-

Table 3: **Panel Regressions of Leverage on Price Flexibility (Errors-in-variables)**

This table reports the results of regressing long-term debt to total assets on the frequency of price adjustment, FPA, and firm characteristics. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. $Prof$ is operating income over total assets, Size is the logarithm of sales, BM is the book-to-market ratio, It2A is intangible assets to total assets, PCM is the price-to-cost margin, and HHI is the Herfindahl-Hirschman index of sales at the Fama & French 48 industry level. Stock-level data are from CRSP and financial statement data are from Compustat. The sample period is January 1982 to December 2014.

	OLS	3rd cum	4th cum	5th cum
	(1)	(2)	(3)	(4)
FPA	0.18∗∗∗	0.26 ∗ ∗	0.21∗∗∗	0.10∗∗∗
	(4.96)	(2.23)	(4.15)	(3.67)
Prof	−0.22∗∗∗	0.18	−0.19 ∗ ∗	−0.31∗∗∗
	(−3.02)	(0.79)	(−2.41)	(−4.23)
Size	0.00	−0.02	0.01 ∗ ∗	0.02∗∗∗
	−1.33	(−1.56)	−2.06	−4.81
BM	0.05∗∗∗	0.09	0.10∗∗∗	0.09∗∗∗
	(5.31)	(1.51)	(4.76)	(7.68)
It2A	0.11∗∗∗	0.66 ∗ ∗	0.03	−0.14∗∗∗
	(3.83)	(2.56)	(0.42)	(−4.31)
PCM	0.00	−0.15∗	0.05	0.09∗∗∗
	(−0.12)	(−1.83)	(1.60)	(3.23)
HHI	−0.03	−0.12∗	0.00	0.02
	(−0.72)	(−1.66)	(0.06)	(0.45)
Constant	0.12∗∗∗	0.13∗	0.04	0.04
	(3.47)	(1.85)	(1.08)	(1.10)
Nobs	8,824			
Adjusted R^2	0.16			

t-stats in parentheses
$*p < 0.10, **p < 0.05, ***p < 0.01$

Table 4: **Triple Differences: Interstate Bank Branching Efficiency Act, Price Flexibility, and Leverage**

This table reports the results for estimating the following linear specification:

$$Lt2A_{i,t} = \alpha + \beta \times FPA_i \times Deregulated_{i,t} + \delta_1 \times FPA_i + \delta_2 \times Deregulated_{i,t} + \eta_t + \eta_f + \epsilon_{i,t},$$

where $Lt2A$ is the long-term debt to assets ratio, FPA is the firm-level frequency of price adjustment, $Deregulated_{i,t}$ is an indicator that equals 1 if firm i is in a state that had implemented the deregulation in year t, and 0 otherwise. η_t and η_f are a full set of year and industry fixed effects. $IndustryFE1$ is a set of eight dummies that capture 1-digit SIC codes. $IndustryFE2$ is a set of forty-eight dummies that capture the Fama & French 48 industries. The sample period is January 1982 to December 2014. Standard errors are clustered at the firm level. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics.

	FPA			FPA Dummy		
	(1)	(2)	(3)	(4)	(5)	(6)
FPA × Deregulated	−0.15***	−0.16***	−0.16***	−0.04**	−0.04**	−0.04***
	(−4.08)	(−4.25)	(−4.64)	(−2.41)	(−2.31)	(−2.71)
FPA	0.30***	0.16***	0.17***	0.09***	0.05***	0.06***
	(8.04)	(4.84)	(5.16)	(5.77)	(3.05)	(3.58)
Deregulated	0.05***	0.02	0.03**	0.04***	0.01	0.01
	(5.75)	(1.50)	(2.15)	(3.39)	(0.86)	(0.78)
Constant	0.15***	0.20***	0.17***	0.16***	0.21***	0.18***
	(19.75)	(8.31)	(8.71)	(15.67)	(7.55)	(6.60)
Year FE		X	X		X	X
Industry FE1		X			X	
Industry FE2			X			X
Nobs	9,119	9,119	9,119	4,558	4,558	4,558
Adjusted R^2	0.08	0.20	0.27	0.09	0.19	0.30

t-stats in parentheses
$*p < 0.10, **p < 0.05, ***p < 0.01$

Table 5: **Triple Differences: Effect at Impact and at alternative horizons**

This table reports the results for estimating the following linear specification:

$$Lt2A_{i,t} = \alpha + \beta \times FPA_i \times Deregulated_{i,t} + \delta_1 \times FPA_i + \delta_2 \times Deregulated_{i,t} + \eta_t + \eta_f + \epsilon_{i,t},$$

where $Lt2A$ is the long-term debt to assets ratio, FPA is the firm-level frequency of price adjustment, $Deregulated_{i,t}$ is an indicator that equals 1 if firm i is in a state that had implemented the deregulation in year t, and 0 otherwise. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. In column (1), the sample only includes firm-level observations in the year before and the year after the implementation of the interstate bank branching deregulation in the state where the firm is headquarterd. In columns (2)-(5), the sample period is indicated on top of each column. Standard errors are clustered at the firm level.

	At Impact	1994-2002	1991-2005	1988-2008	1985-2011
	(1)	(2)	(3)	(4)	(5)
FPA × Deregulated	$-0.07**$	$-0.10**$	$-0.11***$	$-0.12**$	$-0.14**$
	(-2.22)	(-2.37)	(-3.08)	(-3.34)	(-3.81)
FPA	$0.28**$	0..31**	$0.30***$	$0.29***$	$0.29**$
	(5.18)	(6.31)	(6.70)	(6.78)	(7.45)
Deregulated	$0.03***$	$0.04***$	$0.04***$	$0.04***$	$0.04***$
	(4.00)	(4.45)	(4.61)	(4.36)	(5.03)
Constant	$0.17***$	$0.16***$	$0.16***$	$0.16***$	$0.16**$
	(16.24)	(16.64)	(18.24)	(18.82)	(19.35)
Nobs	599	2,795	4,605	6,286	7,857
Adjusted R^2	0.08	0.08	0.08	0.07	0.07

t-stats in parentheses

$*p < 0.10, **p < 0.05, ***p < 0.01$

Table 6: **Triple Differences: Heterogeneous Effect by Dependence on External Financing**

This table reports the results for estimating the following linear specification:

$$Lt2A_{i,t} = \alpha + \beta \times FPA_i \times Deregulated_{i,t} + \delta_1 \times FPA_i + \delta_2 \times Deregulated_{i,t} + \eta_t + \eta_f + \epsilon_{i,t},$$

where Lt2A is the long-term debt to assets ratio, FPA is the firm-level frequency of price adjustment, $Deregulated_{i,t}$ is an indicator that equals 1 if firm i is in a state that had implemented the deregulation in year t, and 0 otherwise. Standard errors are clustered at the firm level. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. The sample period is January 1982 to December 2014.

	FPA		FPA Dummy	
	Low Cash	High Cash	Low Cash	High Cash
	(1)	(2)	(3)	(4)
FPA × Deregulated	−0.19∗∗∗	−0.06	-0.06***	−0.01
	(−4.51)	(−0.92)	(-3.06)	(−0.47)
FPA	0.26∗∗∗	0.14∗∗∗	0.08***	0.04 ∗ ∗
	(7.33)	(2.75)	(5.29)	(2.02)
Deregulated	0.03∗	0.04 ∗ ∗	0.02	0.04
	(1.94)	(2.16)	(0.89)	(1.42)
Constant	0.18∗∗∗	0.08∗∗∗	0.18***	0.11∗∗∗
	(17.47)	(7.54)	(14.51)	(6.01)
Year FE	X	X	X	X
Nobs	6,075	3,044	3,151	1,407
Adjusted R^2	0.08	0.08	0.09	0.07

t-stats in parentheses
$*p < 0.10, **p < 0.05, ***p < 0.01$

Table 7: **Falsification Tests: Early vs. Late Deregulating States**

*This table reports the results of regressing long-term debt to total assets on the frequency of price adjustment, FPA; a dummy that equals 1 for years after 1996, post*1996*; a dummy that equals 1 for firms in states that implemented the interstate bank branching deregulation in the first wave, between 1996 and 1998, early; and all the interactions between these variables. In the first falsification test of columns (4)-(6), a dummy that equals 1 for years after 1992, post*1992*, replaces post*1996*. In columns (1)-(3), the sample period is January 1982 to December 1999; in columns (4)-(6), it is January 1982 to December 1995; in columns (4)-(6), it is January 1982 to December 1995 and January 2001 to December 2014. Standard errors are clustered at the firm level. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. Stock-level data are from CRSP and financial statement data are from Compustat.*

	All (1)	Low Cash (2)	High Cash (3)	Falsification Test 1 (4)	Falsification Test 2 (5)
FPA × post1996 × early	$-0.17**$	$-0.16*$	0.21		-0.01
	(-2.00)	(-1.78)	(0.84)		(-0.09)
FPA × post1996	0.08	0.08	-0.23		$-0.14*$
	(0.99)	(0.95)	(-0.95)		(-1.89)
FPA × early	0.01	0.00	-0.06	0.04	0.01
	(0.16)	(-0.02)	(-0.44)	(0.52)	(0.16)
post1996 × early	0.02	0.00	0.00		0.00
	(0.88)	(0.17)	(-0.09)		(0.12)
FPA	$0.28***$	$0.27***$	0.18	$0.27***$	$0.28***$
	(4.37)	(3.83)	(1.62)	(3.81)	(4.37)
post1996	0.02	0.02	0.02		$0.05**$
	(0.87)	(0.89)	(0.59)		(2.40)
early	0.00	0.03	-0.02	0.00	0.00
	(0.20)	(1.24)	(-0.86)		(0.20)
FPA × post1992 × early				-0.12	
				(-1.39)	
FPA × post1992				0.06	
				(0.73)	
post1992 × early				0.02	
				(0.88)	
post1992				-0.01	
				(-0.48)	
Constant	$0.15***$	$0.16***$	$0.11***$	$0.15***$	$0.15***$
	(7.27)	(6.96)	(7.07)	(6.76)	(7.27)
Nobs	5,376	3,796	1,580	4,110	7,549
Adjusted R^2	0.10	0.10	0.02	0.10	0.08

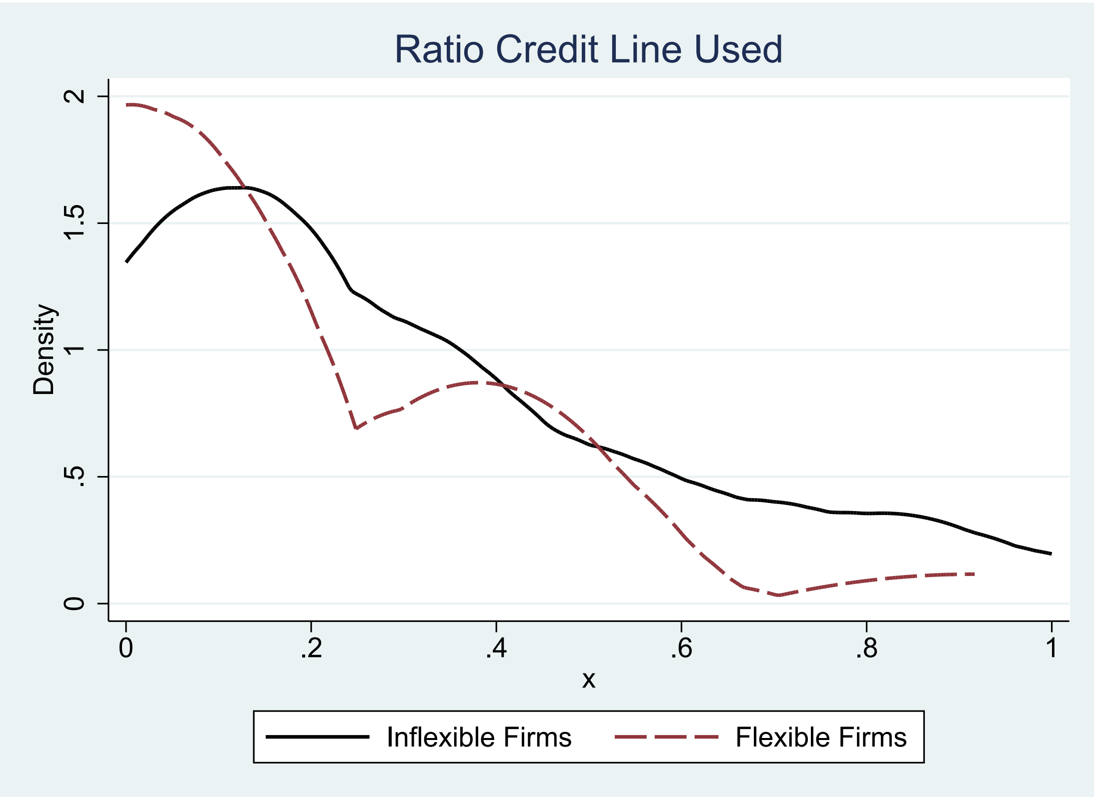
t-stats in parentheses
$*p < 0.10, **p < 0.05, ***p < 0.01$

Online Appendix:
Flexible Prices and Leverage

Francesco D'Acunto, Ryan Liu, Carolin Pflueger and Michael Weber

Not for Publication

Figure A.1: **Intensive Margin of Bank Credit Lines**



This figure plots the density of the share of existing credit lines used by the firms in our sample which also belong to the random sample for which Sufi (2009) collects detailed information on the characteristics of credit lines over time. The solid black is the density for inflexible-price firms. The dashed red line is the density for flexible-price firms. Inflexible-price firms are firms in the bottom 25^{th} *percentile of the distribution by price flexibility. Flexible-price firms are firms in the top* 25^{th} *percentile of the distribution by price flexibility. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. Sample period is January 1980 to December 2013.*

Table A.1: **Price Flexibility and Likelihood of Default**

This table reports the results of logit regressions regressing future defaults on the frequency of price adjustment, FPA, and total debt. Robust standard errors are reported in parentheses. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. Total Debt is the ratio of total debt to sum of liability and market capitalization. The default data are from the Moody's default database. The dependent is a dummy for having a default within years 1 to 5. The sample period is January 1980 to December 2013.

	Def_{t+1}	Def_{t+2}	Def_{t+3}	Def_{t+4}	Def_{t+5}
FPA	-2.02	-2.13*	-1.84*	-1.80**	-1.68**
	(-1.24)	(-1.81)	(-1.91)	(-2.14)	(-2.26)
Total Debt	6.89***	6.16***	5.68***	5.36***	4.93***
	(7.25)	(9.71)	(10.75)	(11.37)	(11.65)
Constant	-7.68***	-6.68***	-6.11***	-5.69***	-5.32***
	(-18.99)	(-25.17)	(-28.02)	(-30.09)	(-32.17)
Observations	13,092	13,092	13,092	13,092	13,092
Pseudo R^2	0.097	0.084	0.075	0.069	0.060

t-stats in parentheses

$*p < 0.10, **p < 0.05, ***p < 0.01$

Table A.2: **Interstate Bank Branching Deregulation, Price Flexibility, and Leverage - Excluding Utilities and Financials**

This table reports the results of regressing long-term debt to total assets on the frequency of price adjustment, FPA; a dummy which equals 1 for years after the state where the firm operates had started to implement the interstate bank branching deregulation, Deregulated; the interaction term between FPA and the dummy, FPA × Deregulated; and firm characteristics. Standard errors are clustered at the firm level. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. $IndustryFE1$ is a set of eight dummies that capture one-digit SIC codes. $IndustryFE2$ is a set of forty-eight dummies that capture the Fama & French 48 industries. The sample period is January 1982 to December 2014.

	FPA			FPA Dummy		
	(1)	(2)	(3)	(4)	(5)	(6)
FPA × Deregulated	−0.15***	−0.15***	−0.15***	−0.04*	−0.04*	−0.04*
	(−3.33)	(−3.44)	(−3.62)	(−1.86)	(−1.75)	(−1.86)
FPA	0.24***	0.20***	0.22***	0.07***	0.05***	0.06***
	(5.89)	(5.07)	(5.26)	(3.86)	(2.96)	(3.52)
Deregulated	0.05***	0.02	0.02	0.04***	0.01	0.00
	(5.92)	(1.02)	(1.57)	(3.23)	(0.44)	(0.17)
Constant	0.15***	0.17***	0.15***	0.16***	0.20***	0.17***
	(19.22)	(7.23)	(7.67)	(16.21)	(6.95)	(6.09)
Year FE		X	X		X	X
Industry FE1		X			X	
Industry FE2			X			X
Observations	7,644	7,644	7,644	3,693	3,693	3,693
Adjusted R^2	0.06	0.09	0.18	0.05	0.11	0.23

t-stats in parentheses
$*p < 0.10, **p < 0.05, ***p < 0.01$

Table A.3: **Panel Regressions of Leverage on Price Flexibility (by industry)**

This table reports the results of regressing long-term debt to total assets on the frequency of price adjustment, FPA; a dummy which equals 1 for years after the state where the firm operates had started to implement the interstate bank branching deregulation, Deregulated; the interaction term between FPA and the dummy, FPA × Deregulated; and firm characteristics. Standard errors are clustered at the firm level. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. $IndustryFE1$ *is a set of eight dummies that capture one-digit SIC codes. The sample period is January 1982 to December 2014.*

	Mining (1)	Food (2)	Materials (3)	Utilities (4)	Trade (5)	Services (6)
FPA × Deregulated	−0.31∗∗	−0.24∗∗∗	0.04	−0.04	−0.21	−0.93∗∗
	(−2.68)	(−3.51)	(0.42)	(−0.42)	(−0.87)	(−2.37)
FPA	0.42∗∗∗	0.23∗∗∗	0.04	−0.12∗	0.54	0.75∗
	(3.71)	(4.47)	(0.73)	(−2.00)	(1.60)	(1.94)
Deregulated	0.05	0.041∗	0.00	0.04	0.02	0.17∗∗∗
	(0.61)	(1.91)	(−0.05)	(1.36)	(0.27)	(4.17)
Constant	0.15∗∗∗	0.17∗∗∗	0.16∗∗∗	0.34∗∗∗	0.16∗∗∗	0.11∗∗∗
	(3.11)	(11.81)	(10.07)	(24.01)	(2.84)	(4.02)
Nobs	405	3,017	3,683	1,275	423	116
Adjusted R^2	0.16	0.08	0.03	0.07	0.14	0.36

t-stats in parentheses
$*p < 0.10, **p < 0.05, ***p < 0.01$

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Table A.4: **Panel Regressions of Leverage on Price Flexibility (Total Debt)**

This table reports the results of regressing total debt to total assets on the frequency of price adjustment, FPA, and firm characteristics. Standard errors are clustered at the firm level. Columns (1) to (3) use the continuous measure of the frequency of price adjustment and columns (4) to (6) use a dummy which equals 1 if the firm is in the top tertile of the frequency of price adjustment distribution. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. $Prof$ is operating income over total assets, $Cf2A$ is income and extraordinary items plus depreciation and amortization to total assets, $C2A$ is cash and short-term investments to total assets, $It2A$ is intangible assets to total assets, $Size$ is log of sales, BM is the book-to-market ratio, PCM is the price-to-cost margin, and HHI is the Herfindahl-Hirschman index of sales at the Fama & French 48 industry level. Stock-level data are from CRSP and financial statement data are from Compustat. The sample period is January 1982 to December 2014.

	FPA			FPA Dummy		
	(1)	(2)	(3)	(4)	(5)	(6)
FPA	0.10∗∗∗	0.03	0.09∗∗∗	0.04 ∗ ∗	0.01	0.03∗
	(2.69)	(0.80)	(2.69)	(2.46)	(0.78)	(1.84)
Prof	−0.27∗∗∗	−0.35∗∗∗	−0.42∗∗∗	−0.22 ∗ ∗	−0.32∗∗∗	−0.35∗∗∗
	(−3.33)	(−4.68)	(−5.86)	(−2.41)	(−3.86)	(−4.02)
Size	0.01∗∗∗	0.02∗∗∗	0.02∗∗∗	0.00	0.01	0.01
	(3.08)	(3.22)	(3.68)	(0.84)	(1.45)	(1.61)
BM	0.05∗∗∗	0.01	0.00	0.06∗∗∗	0.04∗∗∗	0.02∗
	(4.52)	(0.64)	(0.08)	(4.72)	(2.84)	(1.80)
It2A	0.04	0.09∗∗∗	0.11∗∗∗	0.11 ∗ ∗	0.15∗∗∗	0.14∗∗∗
	(1.29)	(2.65)	(3.23)	(2.23)	(3.07)	(3.30)
PCM	0.02	0.01	0.11∗∗∗	−0.02	0.00	0.09 ∗ ∗
	(0.57)	(0.44)	(3.08)	(−0.39)	(0.06)	(2.07)
HHI	0.02	0.12 ∗ ∗	0.14 ∗ ∗	0.05	0.18∗	0.17∗
	(0.28)	(2.06)	(2.20)	(0.61)	(1.90)	(1.84)
Constant	0.12∗∗∗	0.16∗∗∗	0.16∗∗∗	0.18∗∗∗	0.18∗∗∗	0.21∗∗∗
	(3.02)	(2.93)	(2.96)	(3.60)	(2.71)	(3.39)
Year FE1	X	X	X	X	X	X
Industry FE1		X			X	
Industry FE2			X			X
Nobs	8,838	8,838	8,838	4,413	4,413	4,413
Adjusted R^2	0.10	0.20	0.30	0.13	0.22	0.35

t-stats in parentheses

$*p < 0.10, **p < 0.05, ***p < 0.01$

Table A.5: **Interstate Bank Branching Deregulation, Price Flexibility, and Leverage (Total Debt)**

This table reports the results of regressing long-term debt to total assets on the frequency of price adjustment, FPA; a dummy which equals 1 for years after the state where the firm operates had started to implement the interstate bank branching deregulation, Deregulated; the interaction term between FPA and the dummy, FPA $\times$ debranch; and firm characteristics. Standard errors are clustered at the firm level. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. $IndustryFE1$ is a set of eight dummies that capture 1-digit SIC codes. $IndustryFE2$ is a set of forty-eight dummies that capture the Fama & French 48 industries. The sample period is January 1982 to December 2014.

	FPA			FPA Dummy		
	(1)	(2)	(3)	(4)	(5)	(6)
FPA $\times$ Deregulated	$-0.09**$	$-0.09**$	$-0.11***$	-0.03	-0.03	$-0.04*$
	(-2.16)	(-2.14)	(-2.64)	(-1.50)	(-1.35)	(-1.91)
FPA	0.22***	0.11***	0.15***	0.07***	0.04**	0.05***
	(5.38)	(2.61)	(3.66)	(4.33)	(2.00)	(2.74)
Deregulated	0.02	0.01	0.02	0.02	0.00	0.00
	(1.52)	(0.77)	(1.61)	(1.09)	(-0.15)	(-0.03)
Constant	0.24***	0.25***	0.28***	0.24***	0.25***	0.28***
	(26.49)	(9.71)	(9.33)	(20.47)	(8.12)	(7.76)
Year FE		X	X		X	X
Industry FE1		X			X	
Industry FE2			X			X
Nobs	9,133	9,133	9,133	4,563	4,563	4,563
Adjusted R^2	0.03	0.11	0.22	0.05	0.12	0.28

t-stats in parentheses
$*p < 0.10, **p < 0.05, ***p < 0.01$

Table A.6: **Interstate Bank Branching Deregulation, Price Flexibility, and Leverage: Early vs. Late Deregulating States (Total Debt)**

This table reports the results of regressing long-term debt to total assets on the frequency of price adjustment, FPA; a dummy that equals 1 for years after 1996, post1996; a dummy that equals 1 for firms in states that implemented the interstate bank branching deregulation in the first wave, between 1996 and 1998, early; and all the interactions between these variables. Standard errors are clustered at the firm level. Columns (1) to (3) use the continuous measure of the frequency of price adjustment and columns (4) to (6) use a dummy which equals 1 if the firm is in the top tertile of the frequency of price adjustment distribution. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. The sample period is January 1982 to December 2014.

	FPA			FPA Dummy		
	All	Low Cash	High Cash	All	Low Cash	High Cash
	(1)	(2)	(3)	(4)	(5)	(6)
FPA × post1996 × early	−0.13	−0.15∗	−0.11	−0.05	−0.08 ∗ ∗	−0.06∗
	(−1.42)	(−1.74)	(−1.24)	(−1.48)	(−2.12)	(−1.83)
FPA × post1996	0.10	0.13	0.09	0.04	0.07 ∗ ∗	0.06 ∗ ∗
	(1.25)	(1.60)	(1.13)	(1.32)	(2.02)	(2.09)
FPA × early	−0.05	−0.12	−0.13∗	−0.01	0.01	−0.02
	(−0.53)	(−1.35)	(−1.71)	(−0.33)	(0.20)	(−0.62)
post1996 × early	0.00	0.00	−0.01	0.00	0.03	0.02
	(−0.05)	(0.20)	(−0.27)	(−0.24)	(1.07)	(0.99)
FPA	0.25∗∗∗	0.20 ∗ ∗	0.22∗∗∗	0.08 ∗ ∗	0.03	0.05∗
	(3.53)	(2.47)	(3.24)	(2.31)	(1.14)	(1.70)
post1996	0.01	0.00	0.02	0.01	−0.02	−0.01
	(0.51)	(0.17)	(1.07)	(0.72)	(−0.80)	(−0.68)
early	0.03	0.05 ∗ ∗	0.04 ∗ ∗	0.03	0.00	0.00
	(1.34)	(2.05)	(2.00)	(0.84)	(0.43)	(−0.11)
Constant	0.21∗∗∗	0.22∗∗∗	0.21∗∗∗	0.22∗∗∗	0.25∗∗∗	0.255∗∗∗
	(9.97)	(10.89)	(11.72)	(7.70)	(20.94)	(10.57)
Nobs	5,383	5,383	5,383	2,782	2,782	2,782
Adjusted R^2	0.04	0.11	0.20	0.06	0.10	0.28

t-stats in parentheses

$*p < 0.10, **p < 0.05, ***p < 0.01$

Table A.7: **Interstate Bank Branching Deregulation, Price Flexibility, and Leverage (all Firms)**

This table reports the results of regressing long-term debt to total assets on the frequency of price adjustment, FPA; a dummy which equals 1 for years after the state where the firm operates had started to implement the interstate bank branching deregulation, Deregulated; the interaction term between FPA and the dummy, FPA × debranch; and firm characteristics. Standard errors are clustered at the firm level. Equally weighted probabilities of price adjustments are calculated at the firm level using the micro-data underlying the Producer Price Index constructed by the Bureau of Labor Statistics. $IndustryFE1$ *is a set of eight dummies that capture 1-digit SIC codes.* $IndustryFE2$ *is a set of forty-eight dummies that capture the Fama & French 48 industries. The sample period is January 1982 to December 2014.*

	FPA Dummy2		
	(1)	(2)	(3)
FPA × Deregulated	−0.04∗∗∗	−0.04∗∗∗	−0.04∗∗∗
	(−2.62)	(−2.77)	(−3.06)
FPA	0.08∗∗∗	0.04∗∗∗	0.04∗∗∗
	(5.70)	(3.17)	(3.33)
Deregulated	0.04∗∗∗	0.01	0.01
	(5.03)	(0.69)	(1.13)
Constant	0.18∗∗∗	0.22∗∗∗	0.18∗∗∗
	(24.37)	(9.30)	(9.41)
Year FE		X	X
Industry FE1		X	
Industry FE2			X
Nobs	9,119	9,119	9,119
Adjusted R^2	0.05	0.19	0.27

t-stats in parentheses
$*p < 0.10, **p < 0.05, ***p < 0.01$