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. The most flexible-price firms have a 19% higher long-term financial leverage ratio than the most sticky-price firms, controlling for known determinants of capital structure. We rationalize this novel fact in a costly-state-verification model, in which sticky-price firms are more exposed to shocks, and face tighter financial constraints. In the model, a better monitoring ability reduces asymmetric information and narrows the leverage gap between inflexible- and flexible-price firms. Consistently, sticky-price firms increased leverage more than flexible-price firms following the staggered implementation of the Interstate Banking and 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

Firms differ in the frequency with which they adjust output prices to aggregate and idiosyncratic shocks, and these differences are persistent across firms and over time.[1] Firms with rigid output prices are also riskier as they are more exposed to macroeconomic shocks, making price flexibility a viable candidate to explain persistent differences in financial leverage across firms (Lemmon, Roberts, and Zender (2008), DeAngelo and Roll (2015)).

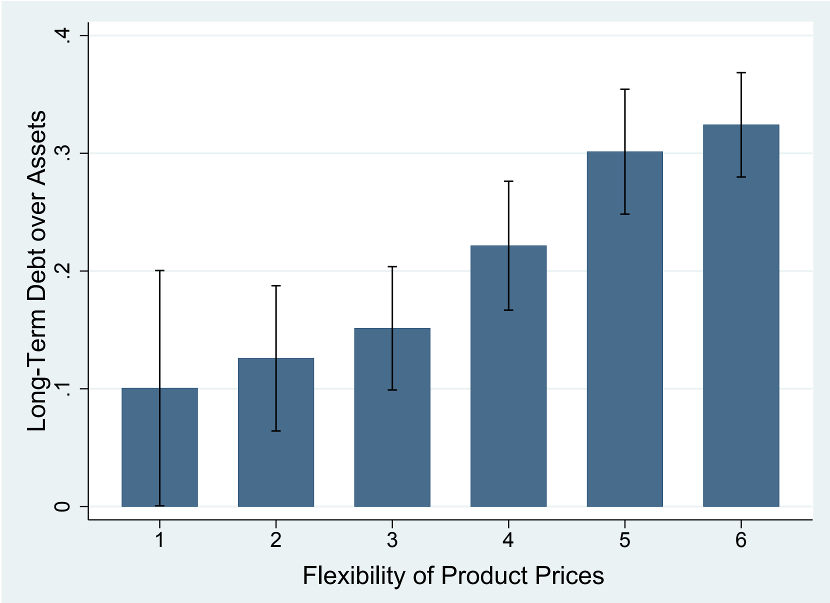
Price rigidity—the fact that firms do not adjust prices to macroeconomic shocks—has long been a focus in Macroeconomics and Industrial Organization. In New Keynesian models, monetary policy has real effects because firms adjust product prices infrequently (Woodford (2003)). Macroeconomic research has studied credit constraints and price rigidity to understand aggregate fluctuations and the effectiveness of monetary policy (Bernanke, Gertler, and Gilchrist (1999)). We provide an empirical link between these two drivers of aggregate fluctuations, and we study their effect on firms' leverage choice.

To guide our empirical analysis, we develop a stylized costly-state-verification model, building on Townsend (1979), Gale and Hellwig (1985), and Bernanke and Gertler (1989). Firms adjust prices imperfectly to shocks, resulting in higher profit volatility for firms with inflexible-product prices. Therefore, the asymmetric information problem is more severe for sticky-price firms when monitoring is costly. Sticky-price firms can credibly pledge a lower share of expected profits ex ante, and have lower leverage compared to flexible-price firms. Increases in the monitoring efficiency of banks allow sticky-price firms to pledge a larger share of expected profits, and hence to increase leverage more than flexible-price firms. The model makes two predictions about the effect of price flexibility on leverage, which we test in the data. First, flexible-price firms have unconditionally higher leverage than sticky-price firms. Second, firms with sticky prices increase leverage more following a positive shock to the effectiveness of monitoring.

Figure 1 documents a novel stylized fact in line with the first model prediction, which is the main result of the paper. We sort firms into six equal-sized groups with increasing

[1] 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 (2016)). Golosov and Lucas (2007) and Alvarez, Gonzales-Rozada, Neumeyer, and Beraja (2011) show that firms' frequency of price adjustment changes little over time, even with inflation rates ranging from 0% to 16%. Weber (2015) documents that price rigidities expose firms more to aggregate risk.

Figure 1: **Flexible Prices and Financial Leverage**



This figure reports the average long-term debt to assets ratio (y-axis) for groups of firms with increasing output price flexibility. We measure the flexibility of product prices at the firm level, using confidential micro data from the Bureau of Labor Statistics (see Section III.A of the paper for a detailed description). For each bin, the graph reports 95% confidence intervals around the mean leverage ratio.

output price flexibility. Moving from firms with the most rigid output prices to firms with the most flexible output prices increases firms' long-term leverage ratio from around 10% to over 30%. We use the confidential micro data underlying the official producer price index (PPI) of the Bureau of Labor Statistics (BLS) to document this fact. We observe monthly good-level pricing data for a subsample of S&P500 firms from January 1982 to December 2014. Empirically, uncertainty about the price level and the role of price-setting frictions is most relevant for profits over long horizons, and we therefore choose long-term leverage as the main outcome variable. Results continue to hold if we look at total leverage or net debt to assets (see Online Appendix).

In the baseline empirical analysis, we find a one-standard-deviation increase in our continuous measure of price flexibility is associated with a 2.1-percentage-point-higher long-term debt-to-assets ratio, which is 10% of the average ratio in the sample (see column (1) of Table 2). We estimate these magnitudes after controlling for size, tangibility, profitability, and the book-to-market ratio. We also control for industry concentration and for firm-level measures of market power and concentration (Hoberg and Phillips (2016)), which might be correlated with the degree of price flexibility of firms because of product-market dynamics.[2] 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)). Results are also similar if we use alternative industry definitions, such as the Fama-French 48 industries, or the Hoberg-Phillips 50 industries (Hoberg and Phillips (2010), Hoberg and Phillips (2016)), which are constructed based on the distance across individual firms in the product space. The size and significance of results are unchanged when we account for measurement error using the errors-in-variables estimator based on linear cumulant equations of Erickson, Jiang, and Whited (2014).

In the model, sticky-price firms are riskier borrowers, because their profits are more volatile. Consistent with this channel, we show firms with inflexible prices are likelier 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 (2016) and Weber (2015)).

We verify that price flexibility is a highly persistent firm characteristic in our sample, consistent with previous findings. 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. This persistence suggests we can hardly consider a shock to firm-level price flexibility for identification purposes in our sample.

This paper does not aim to test for the causal effect of price flexibility on financial leverage, which would require we identify the persistent determinants of the price-setting strategy of firms. Instead, following the model's predictions, we test whether the causal effect of the supply of credit on financial leverage is larger for sticky-price firms than for flexible-price firms. We propose an identification strategy inspired by the financial constraints literature. We (i) identify a positive shock to the supply of bank credit that

[2] Ali, Klasa, and Yeung (2009) show that measures of industry concentration using only publicly-listed firms are weakly correlated with concentration measures using both public and private firms. They find a strong correlation of their Census-based measure with price-to-cost margins. We add both a Compustat-based measure of industry concentration and firm-specific measures of price-to-cost margins.

firms can access, (ii) show sticky-price firms increase leverage more than flexible-price firms after the shock, and (iii) show that the effect does not revert in the short run.

We exploit 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)) as a shock to the availability of bank credit. Restrictions on U.S. banks' geographic expansion date back at least to the 1927 McFadden Act. The IBBEA of 1994 allowed bank holding companies to enter other states and operate branches across state lines, dramatically reshaping the banking landscape in affected states. The step-wise repeal of interstate bank branching restrictions increased the supply of credit. Banking deregulation resulted 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)), more credit cards (Kozak and Sosyura (2015)), more credit lines and subsequent trade credit (Shenoy and Williams (2015)), and increased lending to riskier firms (Neuhann and Saidi (2015)).

We interpret the staggered state-level implementation of the IBBEA as a positive shock to banks' monitoring effectiveness, exogenous to individual firms' financial decisions. This interpretation is consistent with findings in Jayaratne and Strahan (1996), who argue that better quality lending decisions increased economic growth in deregulated states, and Stiroh and Strahan (2003), who find that deregulation led to a reallocation of market shares to better banks and exit of unprofitable banks. It is also consistent with the costly state verification discussion of financial constraints in the literature on aggregate effects of credit constraints (Bernanke and Gertler, 1989; Bernanke, Gertler, and Gilchrist, 1999).[3]

Our empirical design compares 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 sticky-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 required identifying assumptions, we show that before the shock, the trends of long-term debt of flexible- and

[3] An alternative view of banking deregulation proposed by Dick (2006) and Bushman et al. (2016) argues banking deregulation allowed banks to lend to riskier borrowers, possibly due to better geographic diversification. The prediction that riskier sticky-price firms increase leverage more following deregulation than flexible-price firms is robust to such an alternative view of financial constraints.

sticky-price firms are parallel, and that the price flexibility of firms does not change around the shock.

Consistent with the second model prediction, we find sticky-price firms increased leverage more than flexible-price firms after the deregulation. Crucially, sticky-price firms with a lower cash-to-assets ratio, which were more likely to need external financing to fund their operations, drive the effect. The most flexible-price firms kept their leverage virtually unchanged after the deregulation. Results are similar across firms with and without investment-grade bond ratings, alleviating concerns that access to the public bond market drive differences in leverage (see Faulkender and Petersen (2006)).

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

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

In the second falsification test, we use only 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, sticky-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 fear of debt deflation is an important driver of corporate bond yields. Favilukis, Lin, and Zhao (2015) document that firms in industries with higher wage rigidities have higher credit risk. Serfling (2016) finds that more stringent state-level firing laws lower financial leverage of firms headquartered in the state, whereas Simintzi, Vig, and Volpin (2015) show that firms lower their financial leverage in countries passing labor-friendly law changes. Determinants of labor market frictions in this literature vary at the industry, state, or country level, and hence are unlikely to account for our findings, because we exploit variation at the firm level even within industries. In the causal test that exploits the banking deregulation shock, we can also absorb firm-level time-invariant characteristics, such as whether the firms' workforces are unionized or not, and our results do not change.

The paper also speaks to the theoretical and empirical literatures studying the effect of volume flexibility on firms' capital structure. The sign of the effect of volume flexibility on financial leverage is inconclusive. On the empirical side, MacKay (2003) finds that 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 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. The firm's optimal product price is not observable to uninformed lenders, because they cannot observe marginal costs and pricing frictions such as Calvo rates or menu costs. 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 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 credibly to pledge a greater share of real profits to lenders. Bank lending can therefore mitigate the credit constraints which 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:

$$\begin{aligned} Prob\,(\Delta p = 0) &= \pi_0, & (1)\\ Prob\,(\Delta p = h) &= \frac{\pi_h}{2}, & (2)\\ Prob\,(\Delta p = -h) &= \frac{\pi_h}{2}, & (3)\\ \pi_0 + \pi_h &= 1. & (4) \end{aligned}$$

The expected price gap is 0. The parameter h captures how far the firm allows prices to deviate from the optimum when shocks occur either to aggregate or firm-specific demand. 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) show that 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 (2016) calibrate their fully dynamic model to the micro-data underlying the PPI and find similar costs of price adjustments.

In New Keynesian models with monopolistic competition, price dispersion leads to production misallocations and real economic costs (Woodford (2003)). A second-order approximation of the profit function results in an inverted, U-shaped profit function. 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 these features 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 capital K:

$$
\begin{aligned}
Profit_{\Delta p} &= K \times R_{\Delta p}, && (5)\\
R_{\Delta p} &= exp\left(r_{\Delta p}\right), && (6)\\
r_{\Delta p} &= \bar{r} - a(\Delta p)^2. && (7)
\end{aligned}
$$

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. $\bar{r} > 0$ ensures a positive net present value return on capital.[4]

B. The Financing Problem

The owner of the firm has personal wealth or equity, E, which determines the scale of the firm, and has all bargaining power. The lender breaks even in expectation. We normalize the interest rate to zero, and model owner and investors as risk neutral. The total capital

[4]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.

of the firm is the sum of debt, D, 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. The parameter ρ measures the lender's monitoring ability in the model. To ensure that monitoring is always optimal following a bad realization of firm profits, we assume monitoring costs are small relative to the expected gains from monitoring:

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

The revelation principle implies we can focus on an optimal contract, such that the manager never has a reason to lie about the true state of profits. 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:

$$C_0 \quad \geq \quad (1-\rho)K(R_0 - R_h), \tag{14}$$

$$C_0 \quad \leq \quad K\left(R_0 - R_h\right). \tag{15}$$

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

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

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

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

When monitoring is completely ineffective ($\rho = 0$) and free ($\gamma = 0$), equation (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 Producer Price Index (PPI) from the Bureau of Labor Statistics (BLS). We have monthly output 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.[5]

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. 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.

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. Second, it probabilistically selects sample establishments based on the total value of shipments, or on the number of employees, and finally it selects goods within establishments. 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.

We first calculate the frequency of price adjustment (FPA) 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 changes to \$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. This assumption is standard in the literature and does not affect the measure, because sales are rare in the PPI micro data (see Gorodnichenko and Weber (2016)). We then perform two layers of aggregation to create a measure of the frequency of price adjustment at the firm level. We first equally weigh frequencies for all

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

goods of a given establishment using internal identifiers from the BLS.[6] To perform the firm-level aggregation, we manually check whether establishments with the same or similar names are part of the same company. In addition, 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 for all firms in the data set.[7]

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

The price flexibility of similar firms operating in the same industry can differ substantially. 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. (2004)). Assessing the relevance of alternative determinants of price flexibility, which is still a heavily debated topic in Macroeconomics and Industrial Organization, is beyond the scope of this paper. Because our results do not change when we control for firm-level market power and product-market dynamics, firm-level persistent characteristics are likely to determine the within-industry variation in price flexibility across firms.

B. Financial 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.

B.1 Determinants of Financial Leverage

Our preferred measure of leverage, *Lt2A*, is defined as long-term debt over total assets. In the Online Appendix, we show our results are similar if we consider alternative measures of leverage, such as total debt over total assets and net debt over total assets.

All the covariates we use in the analysis are defined at the end the previous fiscal year. To reduce the effects of outliers, we winsorize all variables at the 1$^{\text{st}}$ and 99$^{\text{th}}$ percentiles.

[6]Weighing good-based frequencies by the associated value of shipments does not alter our results.

[7]See Weber (2015) for a more detailed description of the data and the construction of variables. Gorodnichenko and Weber (2016) discuss in detail the number of goods and price spells used to calculate the frequencies at the firm level. The average number of products is 111 and the average number of price spells is 203. See their Table 1.

We follow Lemmon, Roberts, and Zender (2008) and Graham, Leary, and Roberts (2015) in the choice and definition of capital-structure determinants. We define the common determinants of financial leverage as follows: *Profitability* is operating income over total assets, *Size* is the log of sales, *B-M ratio* is the book-to-market ratio, *Intangibility* 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.

B.2 Market Power and Operating Leverage

In the analysis, we also use additional covariates that proxy for market power and operating leverage at the firm level. These controls are important, because the industrial organization literature suggests that product market considerations might affect the price setting strategies of firms. Our preferred measure of market power at the firm level is *Price-Cost margin*, which we define as the ratio of net sales minus the cost of goods sold to net sales. This measure is equivalent to 1 minus operating leverage, and hence it also controls for time-varying changes in operating leverage at the firm level. Our results are unchanged if we control for an alternative measure of operating leverage, the ratio of fixed costs over total sales (see Online Appendix).

To control for industry-level concentration, we use the Herfindahl-Hirschman index (HHI) of annual sales at the Fama-French 48 industry level. Moreover, we use the firm-level definition of concentration within the Hoberg-Phillips industries (*HP Firm-level HHI*), which are constructed based on the distance among firms in the product space, using text analysis to assess the similarity of firms' product descriptions from the annual 10-K filings (see Hoberg and Phillips (2010), Hoberg and Phillips (2016)). These data are available from 1996 onward, which reduces the time span of our analysis. We therefore report the results for the full sample of firm-year observations, and for the restricted sample after 1996 throughout the paper.

B.3 Alternative Definitions of Industries

Product market considerations are likely to be most relevant across industries, as opposed to within industries. In our analysis, we study in detail the effect of within-industry variation in price flexibility on financial leverage, in addition to controlling directly for the proxies for market power described above. A growing literature in finance shows that traditional definitions of industries might not capture the variety of product market spaces in which a firm operates (e.g., see Hoberg and Phillips (2010), Hoberg and Phillips (2016), and Lewellen (2012)). For these reasons, we consider two alternative industry definitions. The first definition is the Fama-French 48 industry taxonomy. The second definition is the Hoberg-Phillips set of 50 industries, based on the distance of firms in their product space (see Hoberg and Phillips (2010), Hoberg and Phillips (2016)).

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. $FPADummy$ 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. The average long-term leverage ratio *Lt2A* is around 21%. Firms have an operative income margin (*Profitability*) of 15%. The average book-to-market ratio is 60% (*B-M ratio*), and the average firm size is USD 3.8 bn. (*Size*). Twenty-one percent of assets are intangible (*Intangibility*). The average price-to-cost margin (*Price-cost margin*) is 37%, and the average industry concentration (*HHI*) is 0.11. Panel B of Table 1 reports the pairwise unconditional correlations among 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. For this reason, in our multivariate analysis we will control for 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, and hence default risk for a given leverage ratio increases. 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 the results for estimating 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 total 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 evidence for defaults adds to previous evidence that sticky-price firms have more volatile profits after shocks and higher unconditional total and idiosyncratic stock return volatility (see Weber (2015) and Gorodnichenko and Weber (2016)).

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 Table A.2 in the Online Appendix, we replicate all the results using total debt and net debt as our measures of leverage.

As a first step, we plot the long-term debt to assets separately for sticky- and flexible-price firms over time. In both panels of Figure 2, 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 dashed lines refer to the ratio of long-term debt over assets of firms in the top quartile by price flexibility, and the black dashed-dotted lines are the differences between the two ratios. In both panels, the baseline prediction of our model seems confirmed: flexible-price firms have on average higher long-term leverage than inflexible-price firms throughout the sample period.

In the top panel of Figure 2, the red vertical line indicates 1996, which is the year the first set of U.S. states started to implement the Interstate Banking and Branching Efficiency Act (IBBEA), an event we describe and exploit for our identification strategy below. In the bottom panel of Figure 2, the red vertical line indicates 2000, which is the year a second group of U.S. states started to implement the IBBEA. In both panels, the difference in the ratio of long-term debt to assets is stable before the deregulation, that is, to the left of the vertical lines, and it declines after the deregulation. 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.

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 absorbs time-varying shocks all firms face, such as changes in economy-wide interest rates; η_k is a set of industry fixed effects, which absorbs time-invariant unobservable characteristics that differ across industries.[8]

The time period varies across specifications because of the availability of the Hoberg-Phillips data. In columns (1) and (5) of Table 2, we consider the full time span of our data from January 1982 until December 2014. In all other columns, the time period is limited from January 1996 to December 2014. This restriction reduces our sample size by

[8]Untabulated results are similar if we limit the variation within industry-years, and hence allow for different trends across industries.

about 50%.[9]

We use two definitions of industry fixed effects. The first definition allows for variation within the 48 Fama-French industries. The second definition follows the 50 industry classification of Hoberg and Phillips (2010) and Hoberg and Phillips (2016). 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.

In columns (1)-(4) of Table 2, FPA is the continuous measure of price flexibility, whereas in columns (5)-(8), 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, as well as measures of market power at the firm level and market concentration at the industry level. Firms with more flexible output prices have a higher ratio of long-term debt to 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 column (2), we add the firm-level measure of concentration within the Hoberg-Phillips industries. The baseline association between the frequency of price adjustment and long-term leverage is virtually unchanged. In columns (3)-(4), we only exploit variation in leverage and the frequency of price adjustment across firms within the same year, and across firms within the same industry. As expected, the size of the association between price flexibility and leverage decreases in the within-industry analysis, because industry-level characteristics are associated with price flexibility. But the baseline association remains economically large and statistically different from 0, which suggests that within-industry variation in price flexibility is an important determinant of capital structure. A t-test for whether the coefficients in columns (3)-(4) differ from the coefficient in column (1) fails to reject the null of no difference at plausible levels of significance.

In columns (5)-(8), we estimate the analogous specifications using the indicator for

[9] Note we cannot restrict the variation within firms, 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.

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) to see whether certain parts of the distribution of the frequency of price adjustment drive our results. This restriction further reduces the sample size, but the results are robust across the alternative sample cuts and we confirm the results we obtained with the continuous measure of price flexibility.[10] Being in the top quarter of the distribution of firms by price flexibility is associated with a six-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 when we add other firm-level controls in equation (18), for instance, when we control for cash over assets, following Faulkender, Flannery, Hankins, and Smith (2012). They 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 allow for measurement error. 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.

In column (1) of Table 3, we report the baseline OLS estimator from column (1) of Table 2 to ease comparison across estimations. In columns (2)–(4), we report the estimated coefficients when implementing the cumulant-equation method of Erickson et al.

[10]The results are similar when we add all other firms and assign them a value of 0 for the FPA dummy measure (see Online Appendix).

(2014) for the third, fourth, and fifth cumulants. We do not report the results for higher-order cumulants because of the sample size. 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 we allow for measurement error in the frequency of price adjustment. 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).

D. Robustness

Table A.2 in the Online Appendix shows that our results do not change when we consider two alternative definitions of financial leverage as our main outcome variable: total debt over total assets and net debt over total assets. Table A.3 in the Online Appendix shows the results are also similar if we use an alternative definition of operating leverage at the firm level, the ratio of fixed costs over total sales. This result suggests persistent determinants of operating leverage can hardly explain our results. In untabulated regressions, we also find that the baseline results are virtually identical when we exclude financial firms and utilities from the sample. Results are also similar when restricting the variation to within industries $\times$ year combinations, both in terms of size and statistical significance. Industry $\times$ year fixed effects control for industry-specific trends in leverage over time.

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.[11] This persistence

[11]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. Therefore, in this paper we do not aim to test for the causal effect of price flexibility on financial leverage.

Instead, based on the predictions of our model, we test whether the causal effect of the supply of debt on financial leverage is higher for sticky-price firms, compared to flexible-price firms. We propose an identification strategy inspired by the financial-constraints literature. We (i) identify a positive shock to the monitoring technology of banks, (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.

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 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. Below, we show the pre-shock trends of long-term leverage for inflexible- and flexible-price firms are similar supporting the parallel-trends assumption. In addition, we do not detect a change in the price flexibility of firms around the shock making it unlikely firms change leverage because their price flexibility changed.

A. Institutional Details and Interpretation

Restrictions to banks' geographic expansion have a long history in the United States (Kroszner and Strahan (2014)). 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.[12]

B. 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

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

credit lines collected by Sufi (2009).[13] These data allow observing 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 the model 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.

C. Triple-Differences Strategy

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

C.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

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

common trends across states 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 3 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 3 plots the estimated coefficients, $\hat{\beta}_t$, and the 95% confidence intervals from the following OLS specification:

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

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 the same state. 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.

C.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 firm-level frequency of price adjustment did not change over time, and the bank deregulation shock did not affect the frequencies. We cannot compute yearly values, because we need several price spells for a given good to construct a meaningful measure.

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 that 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 point estimate is (0.73; 1.12). Also note we truncate price spells by only focusing on a three-year period, and hence we introduce noise into 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.

C.3 Triple-Differences Specification

To implement our strategy, we estimate the following specification:

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

where $Deregulated_{i,t}$ is an indicator that equals 1 if firm i is headquartered in a state that had implemented the deregulation in year t, and 0 otherwise; and η_t and η_I are a full set of year and industry fixed effects. We now can also include a full set of firm fixed effects (η_f), because there is variation in the interaction between price flexibility and deregulation within firms over time. Firm fixed effects absorb industry fixed effects and the frequency of price adjustment. All the results are similar if we also add the full set of controls in equation (18), as well as their interactions with the deregulation dummy.

Equation (20) compares the long-term debt-to-asset ratio 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.

For the purposes of statistical inference, we cluster standard errors at the firm level. All t-statistics are higher if we instead cluster standard errors at the state level, which is the level of the treatment. We only observe firms in 42 different states. The low number of clusters is the likely explanation why standard errors are lower when we cluster at the state level as compared to the firm level.

Table 4 reports the estimates for the coefficients in equation (20). In columns (1)-(4), FPA is the continuous measure of price flexibility; in columns (5)-(8), 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 the 48 industry-level dummies for the Fama-French industry taxonomy. In the third column, we add year fixed effects and the 50 industry-level dummies for the Hoberg-Phillips industry classification. In the fourth column, we add year fixed effects and firm fixed effects.

Across all specifications, the sign of the estimated coefficients are in line with the model predictions. Firms with higher price flexibility have higher long-term debt on average ($\hat{\delta}_1 > 0$). More importantly, across all specifications, we find 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 ($\hat{\beta} + \hat{\delta}_1$) across all specifications. Comparing column (1) with columns (2)-(4), and column (5) with columns (6)-(8), 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. This result survives when we only exploit variation within firms, and hence we absorb any time-invariant determinant of financial leverage at the firm level.

Table A.4 in the Online Appendix shows the results are largely unchanged when we exclude financial firms and utilities. In Table A.5 in the Online Appendix, we find the results for the dummy-variable definition of price flexibility do not change if we consider the full sample of firms, as opposed to restricting the sample to firms in the top 25% and the bottom 25% of the distribution of price flexibility.

C.4 Effect on 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 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. This test aims to estimate the effect of the shock on 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 only have 599 observations compared to 9,119 in our baseline sample. 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 include 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 when we add observations in later years. At the same time, the baseline effect of price flexibility on leverage stays identical across sub-periods. These results are consistent with the idea that it took time for banks to expand across state borders and for firms to adjust their leverage ratios. Diverging trends between flexible- and inflexible-price firms before the shock cannot drive these results, because we find parallel trends before the shock in Figure 3.

C.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. If the deregulation shock is truly driving the interaction effect, then inflexible-price firms that depend more on external finance should drive this effect. We thus estimate the specification in equation (20) separately for firms in the top tercile of cash-to-assets and for other firms. The rationale is that inflexible-price firms with high cash-to-asset ratios will not depend much on external financing. The deregulation shock should instead affect inflexible-price firms with lower cash-to-asset ratios. Consistent with this interpretation, Table 6 shows that the effect of deregulation on firms' leverage is driven by inflexible-price firms with low cash-to-asset ratios (columns (1) and (3)), as opposed to those with high cash-to-asset ratios (columns (2) and (4)).

D. 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 deregulated states.

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

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

Panel A of Figure 4 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-asset ratios. Similar to our earlier results, the subsample of firms with cash-to-asset ratios drive the effects.

We then proceed to assess the validity of our designs by constructing two falsification tests. Panel B of Figure 4 sketches the predictions of our model for the first falsification test. We build on the specification in equation (21), but we limit our estimation to observations before 1996. This limitation implies that no firms, neither in early nor in late states, were 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 level 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 4 sketches the predictions of our model for the 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. We interpret this fact in a costly-state-verification model, in which inflexible-price firms cannot adjust output prices to macroeconomic shocks, and banks can access a monitoring technology. The model also predicts inflexible-price firms should increase their leverage more than flexible-price firms after a positive shock to the monitoring technology of banks. Using the staggered implementation of the 1994 Interstate Bank Branching Efficiency Act across states, we test the second model prediction in a triple-differences identification strategy and find empirical support for it.

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 help us understand the origin of persistent differences in financial leverage across firms as documented 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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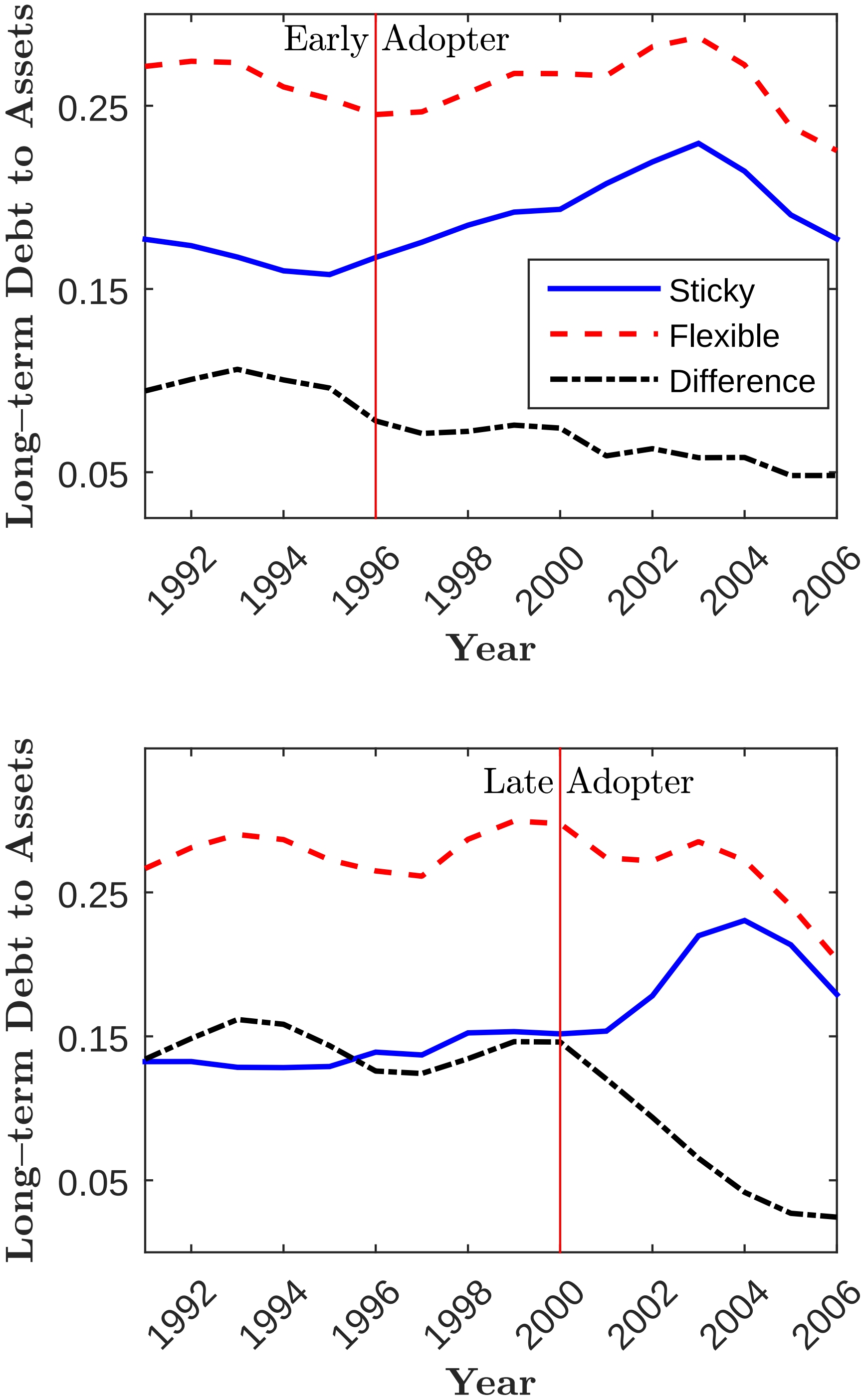
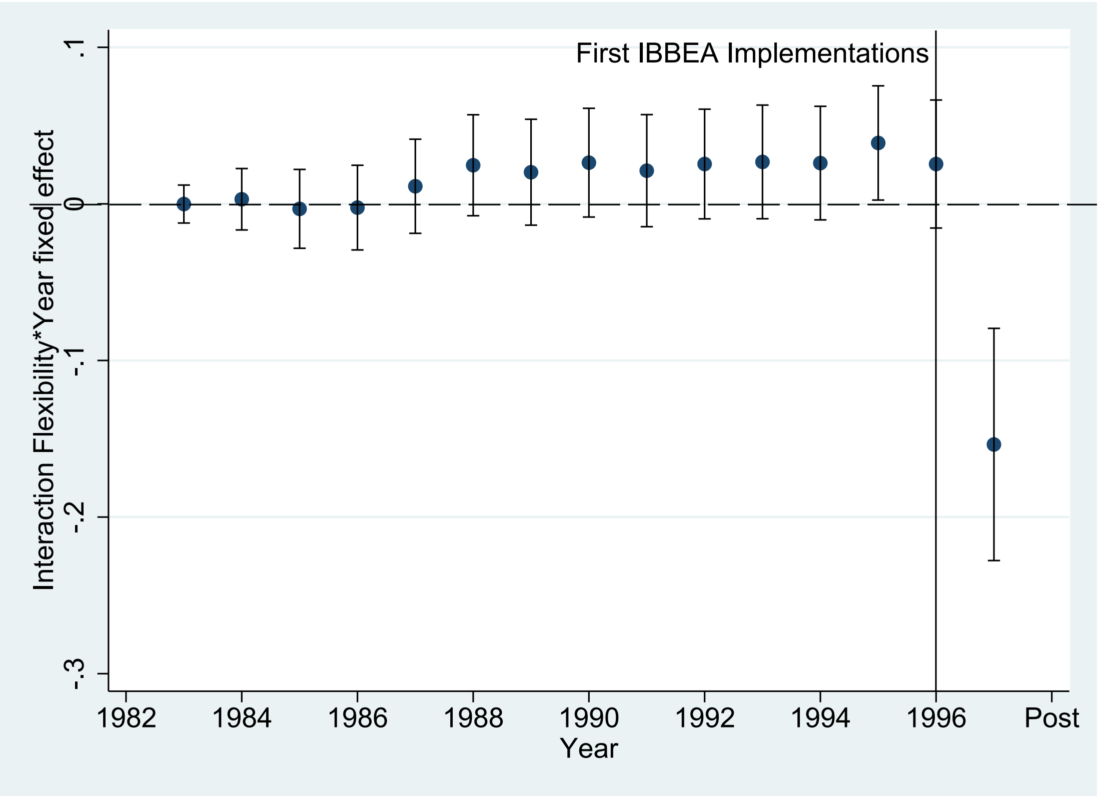
Figure 2: **Long-Term Debt and Price Flexibility**



This figure plots the long-term debt to total assets ratio for different percentiles of the frequency of price adjustment distribution. Sticky-price firms are firms in the bottom quartile of the distribution. Flexible-price firms are firms in the top quartile of the distribution. The sample period is January 1982 to December 2014. 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.

Figure 3: **Parallel Trends Assumption: Assessment of Pre-Trends**

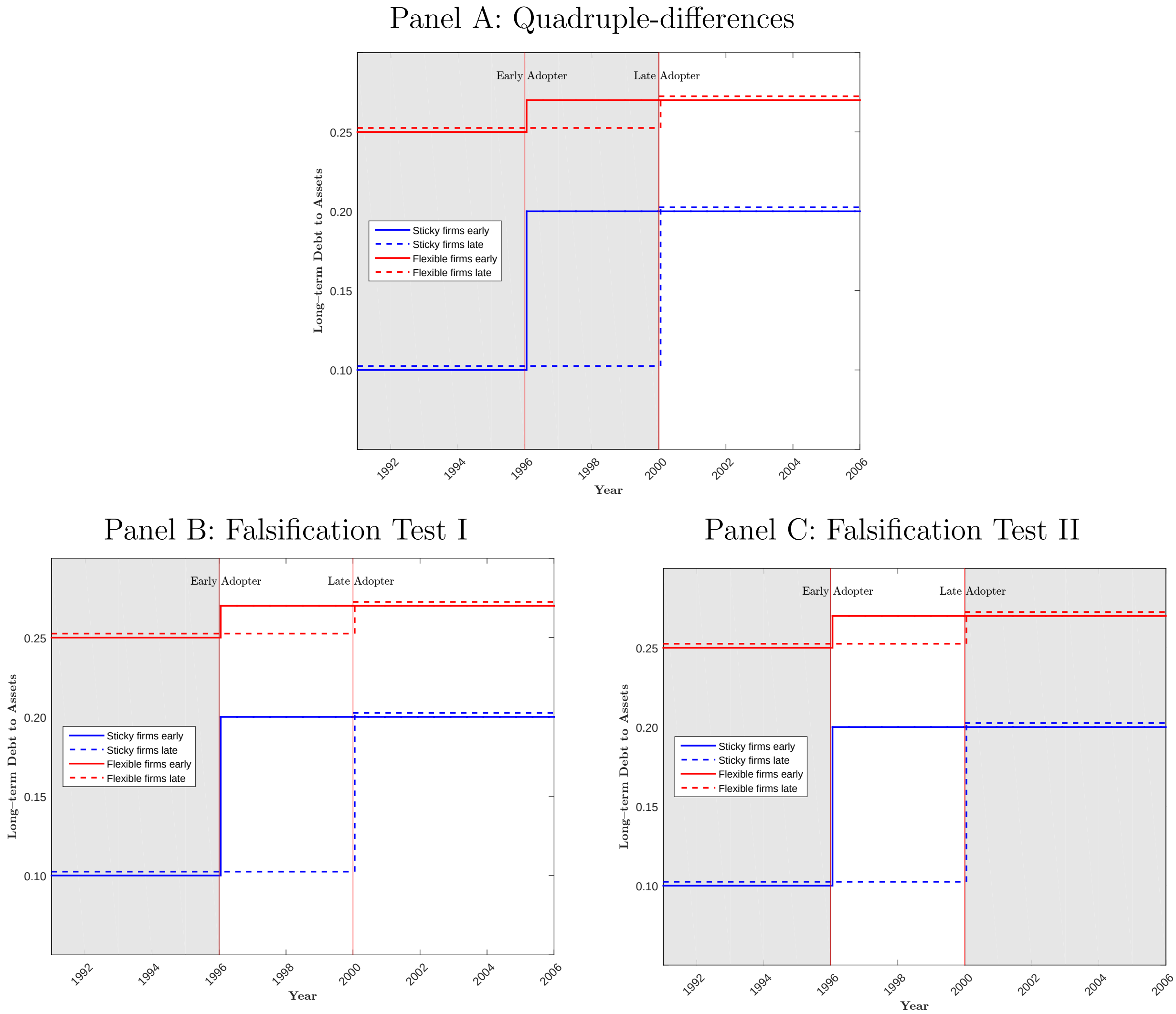


This figure plots the estimated coefficients $\hat{\beta}_t$ *and the 95% confidence intervals from the following linear equation:*

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

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). The sample period is January 1982 to December 2014. Standard errors are clustered at the firm level.*

Figure 4: **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 (blue, dashed). The two top lines refer to flexible-price firms in early states (red, solid) and late states (red, dashed). The model predicts that 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, in which we only use observations up to 1996. Before 1996, no firm was exposed to the deregulation, and hence the model predicts no differences in the 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 inflexible-price firms after the deregulation.

Table 1: **Summary Statistics**

This table reports descriptive statistics for the variables used in the empirical analysis 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. FPA Dummy 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, Profitability is operating income over total assets, Size is the logarithm of sales, B-M ratio is the book-to-market ratio, Intangibility is intangible assets to total assets, Price-cost margin 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***						
Profitability	−0.144***	−0.115***	−0.284***					
Size	0.128***	0.116***	0.121***	−0.0649***				
B-M ratio	0.341***	0.274***	0.258***	−0.456***	−0.0253*			
Intangibility	−0.224***	−0.187***	0.109***	−0.131***	0.297***	−0.192***		
Price-Cost margin	−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 linear 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, Profitability is operating income over total assets, Size is the logarithm of sales, B-M ratio is the book-to-market ratio, Intangibility is intangible assets to total assets, Price-Cost margin is the price-to-cost margin, HHI is the Herfindahl-Hirschman index of sales at the Fama & French 48 industry level, and HP Firm-level HHI is the firm-level measure of product-space concentration based on the Hoberg & Phillips 300 industries. Fama-French 48 FE is a set of forty-eight dummies that capture the Fama & French 48 industries. Hoberg-Phillips 50 FE is a set of fifty dummies that capture the Hoberg & Phllips 50 industries. Stock-level data are from CRSP and financial-statement data are from Compustat. The sample period is January 1982 to December 2014 in columns (1) and (5). The sample is restricted to the period January 1996 to December 2014 in all other columns, due to the availability of the Hoberg-Phillips measures based on textual analysis of 10-K reports. Standard errors are clustered at the firm level. Columns (1) to (4) use the continuous measure of the frequency of price adjustment and columns (5) to (8) use a dummy which equals 1 if the firm is in the top quartile of the frequency of price adjustment distribution and zero if the firm is in the bottom quartile. 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 continuous				FPA dummy			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
FPA	0.18***	0.16***	0.12***	0.08**	0.06***	0.06***	0.04**	0.04**
	(4.96)	(3.60)	(2.98)	(1.98)	(3.93)	(2.89)	(2.11)	(2.29)
Profitability	−0.22***	−0.11	−0.22***	−0.24***	−0.20**	−0.03	−0.12	−0.13
	(−3.02)	(−1.19)	(−2.95)	(−2.92)	(−2.39)	(−0.28)	(−1.23)	(−1.37)
Size	0.00	−0.00	−0.01	−0.01	−0.00	−0.01*	−0.01	−0.01*
	(1.33)	(−0.64)	(−1.16)	(−1.01)	(−0.11)	(−1.80)	(−1.40)	(−1.67)
B-M ratio	0.05***	0.03**	−0.00	0.00	0.06***	0.04***	0.01***	0.01
	(5.31)	(2.57)	(−0.39)	(0.32)	(5.46)	(2.59)	(0.69)	(1.01)
Intangibility	0.11***	0.11***	0.13***	0.09***	0.14***	0.14***	0.16***	0.13***
	(3.83)	(3.08)	(3.76)	(2.79)	(3.23)	(2.67)	(3.39)	(3.14)
Price-Cost margin	0.00	−0.06*	0.04	0.03	0.02	−0.05	0.06	0.06*
	(−0.12)	(−1.81)	(0.99)	(0.83)	(0.40)	(−1.12)	(1.16)	(1.68)
HHI	−0.03	0.05	0.07*	0.01	−0.03	0.03	0.06	−0.15**
	(−0.72)	(0.94)	(1.78)	(0.25)	(−0.48)	(0.38)	(0.90)	(−1.98)
HP Firm-level HHI		−0.04	0.03	0.03		−0.04	−0.02	0.01
		(−1.35)	(1.08)	(1.05)		(−1.19)	(−0.62)	(0.22)
Year FE			X	X			X	X
Fama-French 48 FE			X				X	
Hoberg-Phillips 50 FE				X				X
Nobs	8,824	4,707	4,707	4,672	4,408	2,226	2,266	2,257
Adjusted R^2	0.16	0.09	0.29	0.24	0.18	0.12	0.33	0.32

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 (Lt2A) on the frequency of price adjustment, FPA, and firm characteristics using the linear cumulant equations methodology of Erickson, Jiang, and Whited (2014). Profitability is operating income over total assets, Size is the logarithm of sales, B-M ratio is the book-to-market ratio, Intangibility is intangible assets to total assets, Price-Cost margin 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. Standard errors are clustered at the firm level. All columns use the continuous measure of 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.

	OLS	3rd cum	4th cum	5th cum
	(1)	(2)	(3)	(4)
Frequency Price Adjustment	0.18***	0.26 **	0.21***	0.10***
	(4.96)	(2.23)	(4.15)	(3.67)
Profitability	−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
B-M ratio	0.05***	0.09	0.10***	0.09***
	(5.31)	(1.51)	(4.76)	(7.68)
Intangibility	0.11***	0.66 **	0.03	−0.14***
	(3.83)	(2.56)	(0.42)	(−4.31)
Price-Cost margin	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:

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

where Lt2A is the long-term debt to assets ratio, FPA is the firm-level frequency of price adjustment, and $Deregulated_{i,t}$ 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_k are a full set of year and industry fixed effects. Fama-French 48 FE is a set of forty-eight dummies that capture the Fama & French 48 industries. Hoberg-Phillips 50 FE is a set of fifty dummies that capture the Hoberg & Phllips 50 industries. Firm FE is a set of firm-level fixed effects, which absorbs the measures of price flexibility in column (4) and column (8). The sample period is January 1982 to December 2014 except from coumns (3) and (7), in which the sample period is January 1996 to December 2014, due to the availability of the Hoberg-Phillips measures based on textual analysis of 10-K reports. Standard errors are clustered at the firm level. Columns (1) to (4) use the continuous measure of the frequency of price adjustment and columns (5) to (8) use a dummy which equals 1 if the firm is in the top quartile of the frequency of price adjustment distribution and zero if the firm is in the bottom quartile. 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)	(7)	(8)
FPA × Deregulated	−0.15***	−0.16***	−0.25***	−0.17***	−0.04**	−0.04***	−0.08***	−0.05***
	(−4.08)	(−4.64)	(−4.60)	(−4.56)	(−2.41)	(−2.71)	(−3.30)	(−3.05)
FPA	0.30***	0.17***	0.28***		0.09***	0.06***	0.10***	
	(8.04)	(5.16)	(4.40)		(5.77)	(3.58)	(3.46)	
Deregulated	0.05***	0.03**	0.04**	0.02	0.04***	0.01	0.02	0.02
	(5.75)	(2.15)	(2.42)	(1.43)	(3.39)	(0.78)	(1.15)	(1.45)
Constant	0.15***	0.15***	0.19***	0.18***	0.16***	0.16***	0.15***	0.19***
	(19.75)	(17.84)	(9.91)	(28.36)	(15.67)	(14.42)	(8.94)	(22.20)
Year FE		X	X	X		X	X	X
Fama-French 48 FE		X				X		
Hoberg-Phillips 50 FE			X				X	
Firm FE				X				X
Nobs	9,119	9,119	4,843	9,119	4,558	4,558	2,354	4,558
Adjusted R^2	0.08	0.27	0.21	0.58	0.05	0.19	0.30	0.55

t-stats in parentheses

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

Table 5: **Triple Differences: Effect Before/After and at alternative horizons**

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

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

where Lt2A is the long-term debt to assets ratio, FPA is the firm-level frequency of price adjustment, and $Deregulated_{i,t}$ is an indicator that equals 1 if firm i is in a state that had implemented the deregulation in year t, and 0 otherwise. Stock-level data are from CRSP and financial-statement data are from Compustat. The sample period is January 1982 to December 2014. Standard errors are clustered at the firm level. In column (1), the sample only includes firm-level observations in the year before and after the implementation of the interstate bank branching deregulation in the state where the firm is headquartered. In columns (2)-(5), the sample period is indicated on top of each column. All columns use the continuous measure of 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.

	Before/After (1)	1994-2002 (2)	1991-2005 (3)	1988-2008 (4)	1985-2011 (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:

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

where Lt2A is the long-term debt to assets ratio, FPA is the firm-level frequency of price adjustment, and $Deregulated_{i,t}$ 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 are a full set of year fixed effects. Stock-level data are from CRSP and financial-statement data are from Compustat. The sample period is January 1982 to December 2014. Standard errors are clustered at the firm level. Columns (1) to (4) use the continuous measure of the frequency of price adjustment and columns (5) to (8) 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	
	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 for estimating the following linear specification:

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

where Lt2A is the long-term debt to assets ratio, FPA is the firm-level frequency of price adjustment, and $Deregulated_{i,t}$ *is an indicator that equals 1 if firm i is in a state that had implemented the deregulation in year t, and 0 otherwise.* post1996 *is an indicator that equals 1 in years after 1996. early is an indicator that equals 1 for firms headquartered in states that implemented the interstate bank branching deregulation in the first wave, between 1996 and 1998.* η_t *and* η_k *are a full set of year and industry fixed effects. Stock-level data are from CRSP and financial statement data are from Compustat. The sample period is January 1982 to December 2014. Standard errors are clustered at the firm level. In columns (1)-(3), the sample period is January 1982 to December 1999. In the first falsification test of column (4), an indicator that equals 1 for years after 1992,* post1992*, replaces* post1996*. In column (5), the sample period is January 1982 to December 1995. In column (5), it is January 1982 to December 1995 and January 2001 to December 2014. All columns use the continuous measure of 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.*

	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

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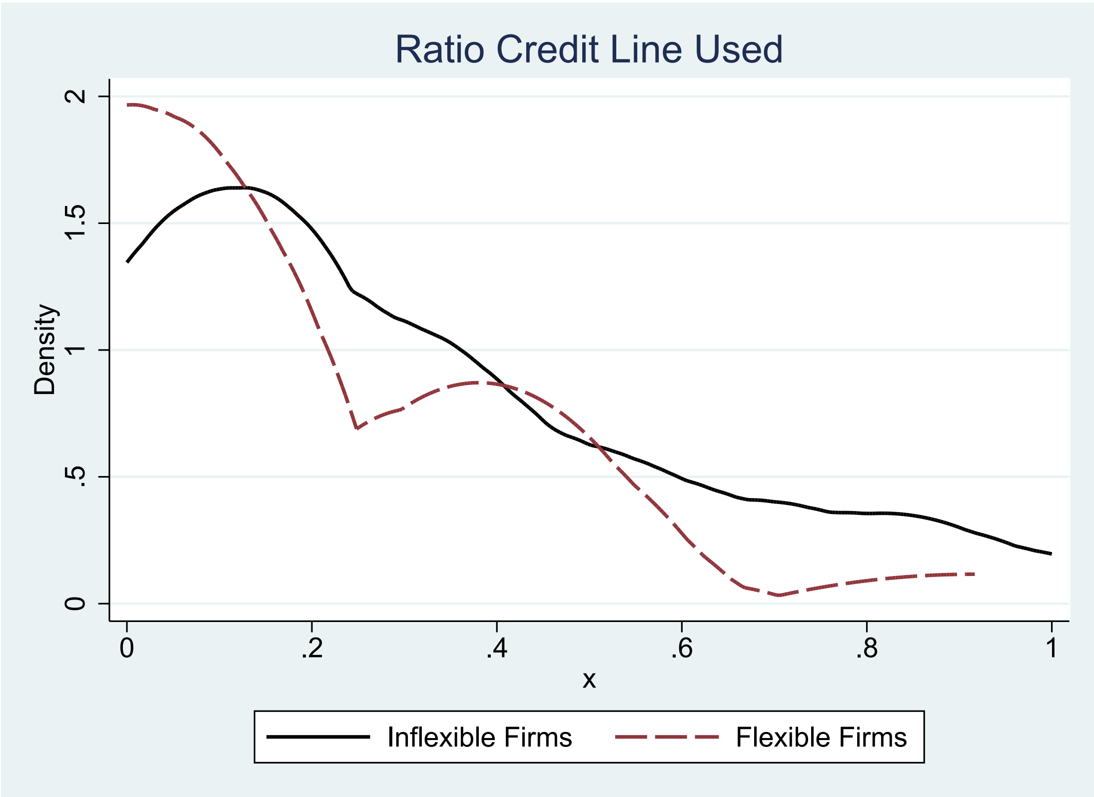
Not for Publication

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



This figure plots the density of the share of existing credit lines used separately for flexible- and inflexible-price firms. The black-solid line is the density for inflexible-price firms. The red-dashed line is the density for flexible-price firms. Inflexible-price firms are firms in the bottom quartile of the frequency of price adjustment distribution. Flexible-price firms are firms in the top quartile of the frequency of price adjustment distribution. The credit line data are from Sufi (2009). The sample period is January 1982 to December 2014. 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.

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 and total debt. Default is a dummy which equals 1 if a firm defaults within the next s years with s running from 1 to 5, FPA is the frequency of price adjustment, and Total Debt is the ratio of total debt to sum of total debt and market capitalization. Default data are from the Moody's default database. The sample period is January 1982 to December 2013. 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.

	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: **Panel Regressions of Leverage on Price Flexibility (Total Debt and Net Debt)**

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

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

where FPA is the frequency of price adjustment, Profitability is operating income over total assets, Size is the logarithm of sales, B-M ratio is the book-to-market ratio, Intangibility is intangible assets to total assets, Price-Cost margin is the price-to-cost margin, HHI is the Herfindahl-Hirschman index of sales at the Fama & French 48 industry level, and HP Firm-level HHI is the firm-level measure of product-space concentration based on the Hoberg & Phillips 300 industries. Fama-French 48 FE is a set of forty-eight dummies that capture the Fama & French 48 industries. Hoberg-Phillips 50 FE is a set of fifty dummies that capture the Hoberg & Phllips 50 industries. Stock-level data are from CRSP and financial-statement data are from Compustat. The sample period is January 1982 to December 2014 in columns (1) and (5). The sample is restricted to the period January 1996 to December 2014 in all other columns, due to the availability of the Hoberg-Phillips measures based on textual analysis of 10-K reports. Standard errors are clustered at the firm level. In columns (1) to (4), Leverage is end-of-year total debt over assets, whereas in columns (5) to (8) it is end-of-year net debt over assets. 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				Net Debt			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
FPA continuous	0.10***	0.12***	0.12***	0.05	0.19***	0.27***	0.17***	0.13**
	(2.69)	(2.48)	(2.87)	(1.23)	(3.17)	(3.18)	(2.65)	(2.01)
Profitability	−0.27***	−0.10	−0.25***	−0.28***	−0.00	0.28**	−0.05	−0.17
	(−3.33)	(−0.92)	(−2.75)	(−3.01)	(−0.02)	(2.04)	(−0.46)	(−1.43)
Size	0.01***	0.01	0.01	0.01	0.03***	0.01	0.02**	0.02*
	(3.08)	(0.80)	(1.03)	(0.90)	(3.60)	(1.29)	(2.10)	(1.81)
B-M ratio	0.05***	0.02	−0.02	−0.01	0.11***	0.07***	0.00	0.01
	(4.52)	(1.62)	(−1.23)	(−0.77)	(6.84)	(3.54)	(0.04)	(0.84)
Intangibility	0.04	0.09**	0.11***	0.08**	0.15***	0.27***	0.36***	0.26***
	(1.29)	(2.19)	(2.75)	(2.06)	(3.07)	(4.20)	(5.72)	(4.84)
Price-Cost margin	0.02	−0.05	0.09*	0.08*	−0.14**	−0.27***	−0.01	0.04
	(0.57)	(−1.46)	(1.88)	(1.75)	(−2.59)	(−4.29)	(−0.19)	(0.61)
HHI	0.02	0.06	0.10*	0.02	−0.03	0.05	0.09	−0.04
	(0.28)	(1.05)	(1.66)	(0.28)	(−0.40)	(0.75)	(1.24)	(−0.71)
HP Firm-level HHI		−0.06*	0.03	0.02		−0.11**	0.05	0.04
		(−1.83)	(0.85)	(0.51)		(−2.09)	(1.13)	(0.95)
Year FE			X	X			X	X
Fama-French 48 FE			X				X	
Hoberg-Phillips 50 FE				X				X
Nobs	8,838	4,715	4,715	4,680	8,838	4,715	4,715	4,680
Adjusted R^2	0.16	0.09	0.29	0.24	0.16	0.15	0.47	0.46

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

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

This table reports the results for estimating the following linear 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, Profitability is operating income over total assets, Size is the logarithm of sales, B-M ratio is the book-to-market ratio, Intangibility is intangible assets to total assets, Fixed Costs to Sales is the firm's operating leverage, HHI is the Herfindahl-Hirschman index of sales at the Fama & French 48 industry level, and HP Firm-level HHI is the firm-level measure of product-space concentration based on the Hoberg & Phillips 300 industries. Fama-French 48 FE is a set of forty-eight dummies that capture the Fama & French 48 industries. Hoberg-Phillips 50 FE is a set of fifty dummies that capture the Hoberg & Phllips 50 industries. Stock-level data are from CRSP and financial-statement data are from Compustat. The sample period is January 1982 to December 2014 in columns (1) and (5). The sample is restricted to the period January 1996 to December 2014 in all other columns, due to the availability of the Hoberg-Phillips measures based on textual analysis of 10-K reports. Standard errors are clustered at the firm level. Columns (1) to (4) use the continuous measure of the frequency of price adjustment and columns (5) to (8) use a dummy which equals 1 if the firm is in the top quartile of the frequency of price adjustment distribution and zero if the firm is in the bottom quartile. 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 continuous				FPA dummy			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
FPA	0.12***	0.11 **	0.15***	0.10 **	0.04 **	0.02	0.04 **	0.05 **
	(3.13)	(2.41)	(3.20)	(2.01)	(2.20)	(1.45)	(2.43)	(2.34)
Profitability	−0.21***	−0.14*	−0.20***	−0.22***	−0.21***	−0.09	−0.11	−0.13
	(−3.62)	(−1.82)	(−2.94)	(−2.90)	(−2.92)	(−0.89)	(−1.36)	(−1.44)
Size	0.00	−0.01	−0.01*	−0.01	−0.00	−0.02 **	−0.02 **	−0.02 **
	(0.28)	(−1.45)	(−1.70)	(−1.55)	(−0.65)	(−2.11)	(−2.07)	(−2.19)
B-M ratio	0.01	−0.00	−0.01*	−0.01	0.02*	0.01 **	−0.00	0.00
	(0.65)	(−0.16)	(−1.70)	(−1.55)	(1.83)	(0.39)	(−0.08)	(0.19)
Intangibility	0.15***	0.14***	0.15***	0.12***	0.17***	0.17***	0.20***	0.15***
	(5.21)	(4.13)	(4.20)	(3.39)	(3.63)	(3.13)	(4.16)	(3.46)
Fixed Costs to Sales	−0.13***	−0.15***	−0.07	−0.10 **	−0.12***	−0.16***	−0.07	−0.06
	(−4.46)	(−4.21)	(−1.53)	(−2.18)	(−2.70)	(−2.72)	(−0.96)	(−0.91)
HHI	0.05	0.08	0.07*	0.03	0.09	0.07	0.05	−0.13*
	(1.00)	(1.57)	(1.79)	(0.57)	(1.47)	(0.96)	(0.87)	(−1.70)
HP Firm-level HHI t		−0.00	0.03	0.03		−0.02	−0.02	0.02
		(−0.08)	(1.10)	(1.09)		(−0.39)	(−0.50)	(0.46)
Year FE			X	X			X	X
Fama-French 48 FE			X				X	
Hoberg-Phillips 50 FE				X				X
Nobs	7,588	4,107	4,107	4,087	3,708	1,931	1,931	1,925
Adjusted R^2	0.14	0.10	0.24	0.20	0.15	0.12	0.32	0.30

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

Table A.4: **Interstate Bank Branching Deregulation, Price Flexibility, and Leverage (Excluding Utilities and Financials)**

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, and $Deregulated_{i,t}$ 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. Fama-French 48 FE is a set of forty-eight dummies that capture the Fama & French 48 industries. Hoberg-Phillips 50 FE is a set of fifty dummies that capture the Hoberg & Phllips 50 industries. Firm FE is a set of firm-level fixed effects, which absorbs the measures of price flexibility in column (4) and column (8). The sample period is January 1982 to December 2014 except from coumns (3) and (7), in which the sample period is January 1996 to December 2014, due to the availability of the Hoberg-Phillips measures based on textual analysis of 10-K reports. Standard errors are clustered at the firm level. Columns (1) to (4) use the continuous measure of the frequency of price adjustment and columns (5) to (8) use a dummy which equals 1 if the firm is in the top quartile of the frequency of price adjustment distribution and zero if the firm is in the bottom quartile. 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)	(7)	(8)
FPA × Deregulated	−0.15***	−0.15***	−0.28***	−0.16***	−0.04*	−0.04*	−0.09***	−0.05**
	(−3.33)	(−3.62)	(−4.57)	(−3.64)	(−1.86)	(−1.86)	(−3.15)	(−2.22)
FPA	0.24***	0.22***	0.36***		0.07***	0.06***	0.11***	
	(5.89)	(5.26)	(5.02)		(3.86)	(3.52)	(3.73)	
Deregulated	0.05***	0.02	0.04**	0.01	0.04***	0.00	0.02	0.01
	(5.92)	(1.57)	(2.13)	(0.55)	(3.23)	(0.17)	(0.99)	(0.77)
Constant	0.15***	0.12***	0.17***	0.15***	0.16***	0.14***	0.20***	0.16***
	(19.22)	(12.88)	(8.34)	(19.94)	(16.21)	(11.95)	(8.71)	(15.14)
Year FE		X	X	X		X	X	X
Fama-French 48 FE		X				X		
Hoberg-Phillips 50 FE			X				X	
Firm FE				X				X
Nobs	7,644	7,644	4,140	7,644	3,693	3,693	1,936	3,693
Adjusted R^2	0.06	0.18	0.17	0.58	0.05	0.23	0.24	0.51

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

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

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, and $Deregulated_{i,t}$ 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. SIC 1-digit FE is a set of eight dummies that capture the 1-digit SIC codes of firms. Fama-French 48 FE is a set of forty-eight dummies that capture the Fama & French 48 industries. Hoberg-Phillips 50 FE is a set of fifty dummies that capture the Hohberg & Phllips 50 industries. Stock-level data are from CRSP and financial-statement data are from Compustat. The sample period is January 1982 to December 2014. Standard errors are clustered at the firm level. All columns use a dummy which equals 1 if the firm is in the top quartile of the frequency of price adjustment distribution and zero 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.

	FPA Dummy2			
	(1)	(2)	(3)	(4)
FPA × Deregulated	−0.04***	−0.04***	−0.04***	−0.06***
	(−2.62)	(−2.77)	(−3.06)	(−2.99)
FPA	0.08***	0.04***	0.04***	0.07***
	(5.70)	(3.17)	(3.33)	(3.16)
Deregulated	0.04***	0.01	0.01	0.02
	(5.03)	(0.69)	(1.13)	(1.27)
Constant	0.18***	0.22***	0.18***	0.21***
	(24.37)	(9.30)	(9.41)	(12.04)
Year FE		X	X	X
SIC 1-digit FE		X		
Fama-French 48 FE			X	
Hoberg-Phillips 50 FE				X
Nobs	9,119	9,119	9,119	4,843
Adjusted R^2	0.05	0.19	0.27	0.21

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