Trade Secrets Protection and Cost Structure

Feng Gao[a] and Xue Wang[b]

[a] *Rutgers Business School, Rutgers University*
[b] *Fisher College of Business, The Ohio State University*

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Abstract

We examine the impact of the risk of rival predation arising from a firm's inability to protect trade secrets on cost elasticity. The empirical setting we use is the staggered adoption of the Inevitable Disclosure Doctrine (IDD) by US state courts over the 1977 to 2011 period, which introduces a plausibly exogenous variation that increases the protection of trade secrets. The results show that the recognition of IDD is associated with an increase in cost elasticity in firms headquartered in IDD recognition states relative to those in non-affected states, and these results are stronger for firms with higher demand uncertainty and greater financial risk, firms facing more intense competition, and firms with a higher ex-ante risk of losing employees to rival firms. Taken together, these results highlight the strategic value of maintaining fixed resources for firms that face greater risk of losing trade secrets to rival firms.

Keywords: cost elasticity; trade secrets; intellectual property; competition

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1. Introduction

One of the fundamental questions in cost accounting is to understand cost behavior and the determinants of cost structure. Prior studies document that cost structure is a function of factors such as asset intensity, demand uncertainty, financial risk, supplier and labor relations (Anderson, Banker and Janakiraman 2003; Banker, Byzalov and Plehn-Dujowich 2014a; Holzhacker, Krishnan and Mahlendorf 2015a; Chen, Kacperczyk, and Ortiz-Molina 2011). While environmental features such as competition and regulation are likely to influence a firm's cost structure, there is surprisingly little research in exploring the role of firms' competitive environment such as the risk of predation by rivals or market entry barrier.

The purpose of our paper is to study the impact of the risk of rival predation arising from a firm's inability to protect trade secrets on cost elasticity. We focus on cost elasticity because it is an important determinant of firm performance. A firm with a more elastic cost structure is less negatively affected by a decrease in demand because a higher proportion of costs will decrease with a reduction in sales (Horngren, Datar and Rajan 2012).

Trade secrets are valuable information existing in all industry sectors, and they provide firms with competitive advantages over their rivals.[1] The inability of firms to protect trade secrets from rival firms could expose them to significant risk of predation, and jeopardize

[1] There are four basic elements that must be present in a trade secret: (1) a trade secret must consist of information; (2) the information must derive economic value (actual or potential) from the fact that it is secret; (3) the information cannot be generally known; (4) the information must be treated as a secret, and be the subject of reasonable efforts to maintain its secrecy (Fenwick and West 2001). Examples of trade secrets include detailed information about a firm's customers, price lists, cost information, information about future business plans, formulas, practices, processes, or designs.

profitability and survival. For example, the US Chamber of Commerce reports in a survey that the theft of trade secrets is associated with more than $50 billion annual loss for firms.[2]

The extent to which firms can protect trade secrets from rivals affects the demand uncertainty and financial risk faced by the firm. Our main hypothesis thus builds on prior work that highlights the strategic value of preserving fixed resource capacity to meet future uncertainty. Greater protection of trade secrets makes it less likely for rival firms to steal customers and gain market shares, thus lowering demand uncertainty and financial risk. Banker et al (2014a) hypothesize and find evidence that firms strategically increase the capacity of fixed resources to reduce the cost of congestion when demand uncertainty increases. Accordingly, we predict that firms are more likely to reduce the capacity of fixed resources and adopt a more elastic cost structure when there is greater protection of trade secrets.

The empirical setting we use to investigate the effect of the risk of losing trade secrets to rivals on a firm's cost elasticity is the staggered adoption of the Inevitable Disclosure Doctrine (IDD) by US state courts over the 1977 to 2011 period. IDD is a legal doctrine through which an employer can claim trade secrets to prevent a former employee from taking a job that may result in the use of trade secrets without the need for proof or evidence.[3] Klasa, Ortiz-Molina, Serfling and Srinivasan (2017) find that the recognition of IDD is significantly associated with reduced mobility of employees in managerial and related occupations to rival firms relative to the mobility of employees in other occupations. This evidence supports the effectiveness of IDD to protect trade secrets from rival firms. However, it is possible that other institutional mechanisms, such as non-competition agreements, might also help restrict outside employment opportunities

[2] "Trends in Proprietary Information Loss," ASIS International, September 2002.
[3] The IDD is applicable even when the employee did not sign a non-compete or non-disclosure agreement with the firm. As an additional analysis, we examine whether the effect of IDD on cost elasticity varies with the extent of enforceability of non-compete agreements.

and prevent rival firms from poaching employees with knowledge of trade secrets. The enforceability of noncompetition agreements also varies significantly across states (Garmaise 2011). Therefore, the empirical effect of IDD adoptions on cost elasticity remains an empirical question.

The staggered adoption of IDD introduces an arguably exogenous variation that enhances the protection of trade secrets. Klasa et al. (2017) document that the state court's decision of adopting IDD is not associated with a state's labor laws, worker characteristics, and economic or political conditions. Therefore, IDD adoptions are unlikely to be systematically related to changes in business or political conditions in a state, lobbying, or to be anticipated by the firm.

We conduct a difference-in-differences analysis to study the impact of trade secrets protection on a firm's cost structure. Trade secrets protection determines a firm's competitive advantages and performance in the product market, so the cost variable we investigate is product cost or cost of goods sold.[4] Following prior literature on cost behavior (Banker et al. 2014a, Halzhacker et al. 2015a), we capture cost elasticity as annual log-changes in cost of goods sold on concurrent annual log-changes in sales revenue.

The key finding is that the recognition of IDD is associated with an increase in cost elasticity in firms headquartered in IDD recognition states relative to those in non-affected states. We show that there are no statistical differences in cost elasticity of treated and control firms in the pre-treatment period, providing support for the use of difference-in-differences analyses to make causal interpretation of the results. Our primary empirical results are robust after including

[4] Selling, General and Administrative Expense (SG&A) is less directly related to the product market relative to cost of goods sold. However, to the extent that it varies with sales volume, we conduct the analysis using SG&A as another cost variable. We find similar statistical significant results as those using cost of goods sold, but with smaller economic magnitudes.

standard economic controls used in cost elasticity tests, controlling for the year, state and firm fixed effects.

We next examine the cross-sectional variation of the impact of IDD on cost elasticity to shed light on the economic mechanisms behind the main finding. We find that the recognition of IDD has a more prominent effect on cost elasticity for firms with higher demand uncertainty and greater financial risk. We also find that it has a less pronounced effect on cost elasticity for firms facing less competitive threats due to higher industry entry barriers and for firms facing a lower ex-ante risk of losing key employees to rival firms. Taken together, these results highlight the strategic value of maintaining fixed resources for firms that face greater risk of losing trade secrets to rival firms.

We conduct an additional analysis to investigate the interaction of IDD and non-competition agreements, another institutional arrangement that restricts outside employment opportunities for employees. The enforcement of these agreements differs significantly across states, which likely creates variation in employment opportunities in rival firms. The results show that the impact of IDD on cost elasticity is weaker in states with stronger enforceability of non-competitive agreements, suggesting that these two mechanisms to restrict outside employment opportunities appear to be substitutes.

Our paper makes three primary contributions to the literature. First, we present new evidence on the impact of competitive market conditions on cost management. A central argument in cost management is how firms respond to exogenous shocks in the environment. The empirical setting used in this paper, the adoptions of IDD, provides an exogenous shock that increases the protection of trade secrets and reduces the competitive risk in the market. Therefore, the results documented in the paper should be more reflective of managers' reactions to outside environmental shocks rather than the mechanical results of cost behavior (Andersen and Lanen 2007).

Second, we contribute to the literature on the benefits and costs of trade secrets protection. Almeling (2012) show that firms are increasingly concerned with trade secrets protection because trade secrets are critical resources and rival firms have strong incentives to gain access to these secrets. Prior research documents positive market reactions to the state courts' adoptions of IDD (Klasa et al 2017; Qiu and Wang 2017). Our paper adds to this line of research by demonstrating that increased elasticity of cost structure might be one channel through which greater protection of trade secrets can enhance firm value.

Finally, we add to the literature on the interactions of firms' operating and financial leverages. Kahl, Lunn and Nilsson (2014) find that firms with higher fixed costs (i.e. higher operating leverage or low cost elasticity) tend to have lower financial leverage and larger cash holdings relative to firms with lower fixed costs, and interpret the evidence as suggesting operating leverage is an important determinant of financial leverage. However, financial risk is also considered an important factor that influences cost elasticity. Holzhacker et al (2015a) document that firms react to heightened financial risk by taking actions to increase cost elasticity. One challenge associated with the above studies is that both operating and financing leverage decisions are endogenous. Together with Klasa et al (2017), we demonstrate that firms might adjust operating and financial leverages simultaneously in response to a common shock in the environment.

The remainder of the paper is organized as follows. Section 2 discusses the IDD background, develops hypothesis and reviews related literature. Section 3 describes our data and sample selection, and section 4 presents the empirical analyses and results. Section 5 concludes the paper.

2. Hypothesis Development

2.1 The Inevitable Disclosure Doctrine

The protection of trade secrets is in general subject to the jurisdiction of state laws.[5] Historically trade secrets laws were based on the precedent state court case decisions until 1979 when the National Conference of Commissioners on Uniform State Laws recommended the enactment of the Uniform Trade Secrets Act (UTSA) in all states. By 2014, 47 states, Washington DC, and the US Virgin Island have enacted UTSA.[6] Importantly, Section 2 of the UTSA allows the owner of a trade secret to obtain injunctive relief to present misappropriation of trade secrets. Misappropriation can be categorized into three types of prohibited actions: (1) wrongful acquisition; (2) wrongful use; and (3) wrongful disclosure of someone else's trade secrets.[7]

The notion underlying the Inevitable Disclosure Doctrine (IDD) is that, when the probability that an employee would reveal trade secrets is high, a court may enjoin the employment to prevent the potential disclosure of information. Thus, the IDD enables a court to find that a former employee would disclose proprietary information in the position with a new employer, even when there is no evidence of actual disclosure. This allows companies seeking to protect trade secrets to assert that the employee would be employed in such a capacity that she

[5] The discussion in this section is primarily based on Fenwick and West (2001), Klasa et al (2017), and Li et al (2017).
[6] The states that have not adopted UTSA include Massachusetts, North Carolina, and New York.
[7] Section 3426.1(b) of the UTSA provides a statutory definition of "misappropriation." Misappropriation does not need to be deliberate; it can occur through negligence or mistake.

would "inevitably" disclose trade secrets. Therefore, the recognition of IDD by a state court greatly enhances the protection of trade secrets for firms located in the state by reducing the risk that departing employees will disclose or use trade secrets in other companies.[8]

2.2 The Impact of IDD Recognitions on Cost Elasticity

Trade secrets are important intellectual property assets for a firm. Unlike patents, copyrights or trademarks, trade secrets are not publicly recognized or registered with the government. However, the lack of protection of trade secrets could be very costly. Once trade secrets are possessed by a rival company, they can be put to immediate use, which might threaten the profitability and survival of the company. For example, if a trade secret involves cost-saving procedures for a manufacturing process, they could be implemented by the rival company without getting caught.

The extent to which firms can protect trade secrets from rivals affects the demand uncertainty and financial risk faced by the firm. Our main prediction about the impact of the IDD on cost elasticity is based on prior work that highlights the strategic value of preserving fixed resource capacity to meet future uncertainty. The recognition of IDD provides greater protection of trade secrets against rival predation, which makes it less likely for rival firms to steal customers and gain market share, thus lowering demand uncertainty and financial risk. Demand uncertainty has an impact on a firm's commitments of "fixed" activity resources. Banker et al (2014a) analytically and empirically study the relationship between demand uncertainty and cost elasticity. The intuition underlying Banker et al (2014a) is that both

[8] Importantly, lawsuits related to employment contracts are filed under the context of employment law. Thus, the relevant jurisdiction for a lawsuit under the IDD is the state where the former employee worked. As such, the protection of trade secrets under the IDD is effective even when the new employer of the departing employee is located in another state that does not adopt the IDD.

unusually high and unusually low demand are possible when demand uncertainty is higher. Thus firms should increase the capacity of fixed resources to reduce the cost of congestion (under the circumstance of unusually high demand) when demand uncertainty increases. Applying the argument to our empirical setting, we predict that firms are more likely to reduce the capacity of fixed resources and adopt a more elastic cost structure after IDD adoptions. Further, greater protection of trade secrets lowers the financial and competitive risk faced by the firm. As a result, firms are more likely to increase the elasticity of the cost structure to increase profits.

While IDD enhances trade secrets protection in states that recognize it, it is possible that other institutional mechanisms, such as non-competition agreements, might also help restrict outside employment opportunities and prevent rival firms from poaching. Therefore, the average effect of IDD adoptions on cost elasticity remains an empirical question.

Next, we develop hypothesis on the cross-sectional variation of the impact of IDD recognitions on cost elasticity. The conceptual framework behind our cross-sectional analyses hinges on the strategic value of maintaining fixed resources for firms that face greater risk of losing trade secrets to rival firms. First, the fixed resource capacity is particularly valuable for firms with higher demand uncertainty and financial risk ex ante. To the extent that the recognitions of IDD reduce demand uncertainty and financial risk, we predict that the effects on cost elasticity should be more prominent for firms operating in industries with greater demand uncertainty and higher financial risk ex ante.

Likewise, the fixed resource capacity is also important for firms with higher competitive risk. We reason that the competitive risk is higher for firms with a greater ex-ante risk of losing employees with knowledge of trade secrets to rivals, and for firms in industries with lower entry barriers and less differentiated products. To the extent that the recognitions of IDD reduce competitive risk, we predict that the effects on cost elasticity should be stronger for firms with a greater ex-ante risk of losing employees with trade secrets knowledge and for firms in industries with lower entry barriers and less differentiated products.

2.3 Related Research

There is a large literature on cost behavior and the determinants of cost structure. The two primary cost structure variables are cost elasticity, defined as the percentage change in cost for each percentage change in quantity, and cost asymmetry (or cost stickiness), defined as the greater increase in cost when there is an increase in quantity relative to the decrease in cost when there is a decrease in quantity (Anderson et al 2003).[9] Our paper is more closely related to the literature on cost elasticity.

Lower cost elasticity imposes more risk on firms because it increases the likelihood of incurring losses that might lead to default on critical obligations and operation disruptions (Horngren et al 2012). Therefore, it is important to gain an understanding of the factors that determine cost elasticity.

To the best of our knowledge, there are three streams of research on the determinants of cost elasticity. The first stream of research focuses on two primary risk drivers: demand uncertainty and financial risk. Banker et al (2014a) hypothesize that firms will choose a higher fixed capacity to reduce congestion costs when uncertainty increases. Using data from

[9] Motivated by the finding in Anderson et al (2013) that the absolute change in SG&A cost associated with decreased sales is significantly less than that associated with increased sales, many papers examine the factors that determine cost asymmetry, such as capacity utilization (Balakrishnan, Petersen and Soderstrom 2004), incentives to manage earnings (Dierynck, Landsman and Renders 2012, Kama and Weiss 2013), the pattern of sales changes (Banker, Byzalov, Ciftci and Mashruwala 2014b), agency costs (Chen, Lu and Sougiannis 2012), adjustment costs such as employee firing costs (Banker, Byzalov and Chen 2013), and litigation risk (Li, Monroe, and Coulton 2017). However the sticky cost model has also been questioned by scholars in recent years (see, for example, Anderson and Lanen 2007, Balakrishnan, Labro and Soderstrom 2014).

manufacturing firms, they find strong empirical results that firms facing higher demand uncertainty have a cost structure with higher fixed and lower variable costs, i.e. lower cost elasticity.

Holzhacker et al (2015a) additionally propose financial risk as another factor that influences cost elasticity. Using a rich dataset from the California hospital industry, they find that both demand uncertainty and financial risk affect cost elasticity and that these effects are moderated by hospitals' resource procurement choices.

The second stream of this research studies the impact of labor and supplier relations on cost structure. Chen et al. (2011) argue that labor unions impose constraints on firms' operating flexibility. Consistent with their conjecture, they find that labor unions are positively associated with operating leverage and cost of equity. Likewise, Serfling (2016) documents that a firm's operating leverage increases following the adoption of state-level labor protection laws, which proxies for an increase in employee firing costs. Focusing on the supply chain, a recent paper by Chang, Hall and Paz (2017) documents that suppliers with greater customer concentration tend to make relationship-specific investments with more rigid cost structures (more fixed-to-variable costs or lower cost elasticity).

The third stream of research investigates the impact of regulations on cost elasticity. Kallapur and Eldenburg (2005) document that hospitals increase cost elasticity in response to fixed-price regulation. Holzhacker, Krishnan and Mahlendorf (2015b) further study the role of ownership on firms' cost elasticity responses to price regulation using data from the German hospital industry.

Our study builds on the prior literature on cost elasticity, and explores the role of firms' competitive environment in cost structures, which only receives limited research attention.[10] A firm's operating and product strategies respond to the competitive environment, and it is intuitive that competitive risk is a direct contributing factor of cost structure. The staggered recognition of IDD provides an arguably exogenous shock that increases the protection of trade secrets and reduces the competitive risk in the market. We are able to use this quasi-experimental setting to provide causal evidence about the impact of trade secrets protection on cost elasticity.

Our paper is also related to the emerging literature on the economic effects of trade secrets protection, with most of these studies using the staggered recognition of IDD as the empirical setting. Overall the recognition (rejection) of IDD has differential effects on the mobility of university-educated workers (or managerial and related occupations) versus other workers (Png and Samila 2015; Klasa et al 2017), suggesting that IDD reduces the mobility of employees that have knowledge of trade secrets. Along those lines, Chen, Gao and Ma (2017) document that IDD recognitions are associated with increased human capital driven acquisitions. The above evidence supports that IDD recognitions help protect trade secrets from predation by the rival firms.

The overall market reactions to the state courts' adoptions of IDD (Klasa et al 2017; Qiu and Wang 2017) are positive and significant, suggesting that trade secrets protection is value enhancing for shareholders. The inferences are more varied in studies that examine the impact of IDD recognitions on firm policies. Some studies argue that IDD recognitions mitigate the risk of losing trade secrets to rivals and reduces uncertainty, and therefore are associated with higher

[10] The only other study we are aware of in exploring the role of competitive environment is Li and Zheng (2016). They use two firm-level text based product competition measures to study the impact of product market competition on cost stickiness, and find that cost stickiness asymmetry increases in product market competition.

financial leverage and higher quality disclosure (Klasa et al 2017; Lin, Wei and Wu 2016). Other papers focus on manager's career concerns, arguing that the adoptions of IDD limit managers' outside opportunities and exacerbate their risk aversion aptitude, resulting in less risk taking in financing and investing activities and greater asymmetric withholding of bad news relative to good news (Chen, Jung, Peng and Zhang 2017; Ali, Li and Zhang 2015).[11]

Our study adds to this line of research by demonstrating an economic channel of the impact of greater trade secrets protection, which is different from prior studies that focus on the financing and disclosure policies. Our results are consistent with the idea that managers' responses in operating decisions to increase the elasticity of cost structure might be one channel through which greater protection of trade secrets can enhance firm value.

3. Sample Selection and Descriptive Statistics

Given that the relevant jurisdiction for a lawsuit under IDD is the state where the former employee worked, we follow Klasa et al. (2017) and start with the sample of firms with headquarters in the U.S. on the merged Compustat-CRSP data between 1977 and 2011. The initial year is chosen to be five years prior to the adoption of IDD in Pennsylvania in 1982, while the last year is five years after the most recent adoption of IDD in Kansas in 2006. After dropping utilities and financial firms, we also require non-missing COGS and sales revenue in two consecutive years. As in Banker et al. (2014a), we delete 1% extreme observations for these two variables on each tail. Our main sample contains a total of 102, 409 firm-year observations. The detail of the sample selection process is reported in Table 1.

[11] Focusing on the role of proprietary costs, Li et al (2017) find that firms reduce the level of disclosure regarding their customers' identities following the IDD adoptions, with the effects concentrated in firms operating in industries with a higher degree of entry threats and industries with a higher degree of volatility.

Table 2 summarizes the variables used in the empirical analysis. Both the mean changes in logarithms of COGS and sales revenue are around 10%. About 42% of the observations are firm-year observations after the adoption of IDD. The annual growth in GDP (in 2009 dollars) is obtained from the Bureau of Economic Analysis. The average GDP growth in the U.S. between 1977 and 2011 is 3%.

4. Empirical Results

4.1 The Impact of the Adoption of IDD on Cost Elasticity

Following Klasa et al. (2017), we use the incidence of a precedent-setting case, which a state would follow in the future, to identify whether IDD is adopted in the state. With an assumption that these events are exogenous, we could evaluate the impact of the adoption of IDD on a firm's cost elasticity using the following difference-in-differences model (Banker et al. 2014a).

$$\Delta\, lnCOGS = \beta_0 + \beta_1\, \Delta\, lnSales + \beta_2\, IDD + \beta_3\, IDD{*}\Delta\, lnSales$$
$$+ \beta_4\, GDP\ growth + \beta_5\, GDP\ growth{*}\ \Delta\, lnSales + FE + \varepsilon. \quad (1)$$

The dependent variable is the change in the natural logarithm of costs of goods sold (COGS) between year *t* and year *t*-1 ($\Delta\, lnCOGS$). The independent variable is the change in the natural logarithm of sales revenue between year *t* and year *t*-1 ($\Delta\, lnSales$). The specification allows for an easy interpretation of the coefficients. The coefficient β_1 measures the elasticity of cost, i.e. the percentage change in COGS with a 1% change in sales revenue, for firms headquartered in a state that has not adopted IDD as of year *t*. The indicator variable, *IDD*, equals to one if a firm's headquarter is in an adopting state in the year of and after the adoption of precedent-setting cases as in Table 1 of Klasa et al. (2017), and zero otherwise. As Compustat only provides a firm's current headquarter state, we follow Gormley and Matsa (2016) and use the data compiled by Cohen (2012) for the historical headquarter state. The data ends in 2006. We fill in missing headquarter information with the headquarter information from Compustat. The coefficient β_3 measures the change in cost elasticity following the adoption of IDD, and we predict $\beta_3 > 0$.

The annual growth in GDP (in 2009 dollars) is measured as the percentage change in GDP between year *t* and *t*-1. It is included to control for the potential variation in cost behavior across different macroeconomic environments. We check the robustness of the results using different types of fixed effects. All standard errors are clustered at the firm level.[12]

Table 3 reports the estimation results of model (1). We include year and industry fixed effects in column (1), year, industry and state fixed effects in column (2), and add firm fixed effects in Column (3). Across all columns, the coefficient on $\Delta\, lnSales$ is positive and significant at the 1% level. As predicted, the coefficient on $IDD{*}\Delta\, lnSales$ is positive and statistically significant at the 1% level, suggesting that the adoption of IDD increases the cost elasticity from 91% to 94%.

The validity of using the difference-in-differences model depends on the assumption of "parallel trends," which means the average change in the outcome variable would have been the same for both the treatment and control groups in the absence of treatment (Roberts and Whited 2013). While the parallel trend assumption is fundamentally untestable, we follow the suggestion in Roberts and Whited (2013) to perform sensitivity test to check the validity of the difference-in-differences tests. Specifically, we examine the timing of the changes in cost elasticity, relative to the timing of the adoption of IDD. We include the following indicator variables: (1) IDD_0 is equal to one if a firm's headquarter is in a state that adopts the IDD in the current year; (2) IDD_{-1}

[12] Results are similar when we cluster the standard errors at the state level.

and IDD_{-2} are equal to one if a firm's headquarter is in a state that adopted the one year ago and two year ago; (3) IDD_1 and IDD_2 are equal to one if a firm's headquarter is in a state that will adopt the IDD in one year and in two years. Our empirical setting is interesting in that it also includes rejections of IDD that was adopted earlier in three states (FL, MI and TX). The indicator variable IDD_r is equal to one for these three states starting from the first year of rejection, and zero otherwise. Each of the indicator variables are then interacted with the change in sales revenue.

$$\begin{aligned}\Delta\ lnCOGS &= \beta_0 + \beta_1\,\Delta lnSales + \beta_2\,IDD_{-2} + \beta_3\,IDD_{-2}{*}\Delta lnSales + \beta_4\,IDD_{-1}\\ &+ \beta_5\,IDD_{-1}{*}\Delta lnSales + \beta_6\,IDD_0 + \beta_7\,IDD_0{*}\Delta lnSales + \beta_8\,IDD_1 + \beta_9\,IDD_1{*}\Delta lnSales\\ &+ \beta_{10}\,IDD_2 + \beta_{11}\,IDD_2{*}\Delta lnSales + \beta_{12}\,IDD_r + \beta_{13}\,IDD_r{*}\Delta lnSales + \beta_{14}\,GDP\ growth\\ &+ \beta_{15}\,GDP\ growth{*}\ \Delta lnSales + FE + \varepsilon. \qquad (2)\end{aligned}$$

A significant coefficient on $IDD_{-1}{*}\Delta lnSales$ or $IDD_{-2}{*}\Delta lnSales$ would suggest potential violations of the parallel trend between the treatment firms (firms in states adopting IDD) and control firms (firms in non-adopting states) in the absence of the event, i.e. the cost elasticity between the treatment and control firms differs even prior to the adoption of IDD. In comparison, a significant coefficient on $IDD_0{*}\Delta lnSales$, $IDD_1{*}\Delta lnSales$ or $IDD_2{*}\Delta lnSales$ could provide evidence of the timing of the impact of IDD on cost elasticity.

Table 4 reports the estimation of model (2) across several specifications of fixed effects. In all specifications, the coefficients on $IDD_{-2}{*}\Delta lnSales$ and $IDD\text{-}_1{*}\Delta lnSales$ are statistically insignificant. However, the coefficients on $IDD_0{*}\Delta lnSales$ are positive and statistically significant at the 5% or the 10% level. The result suggests that there does not seem to be any significant difference in cost elasticity between the treatment and control firms in the pre-IDD adoption period, and the difference only shows up when the IDD is adopted. Hence, reverse causality or a violation of the parallel trends assumption does not seem to drive our main findings.

Our hypothesis about an increase in cost elasticity following the adoption of IDD could be applied to the rejection of IDD with the opposite prediction. Consistent with this prediction, the coefficients on $IDD_r*\Delta\, lnSales$ are negative and statistically significant at the 1% level. The total effect of IDD adoptions is captured by $IDD_0*\Delta\, lnSales$ + $IDD_1*\Delta\, lnSales$+ $IDD_2*\Delta lnSales$, which is positive and significant (p=0.015). The overall effect of adopting IDD first and then reversing it can be measured by $IDD_0*\Delta\, lnSales$ + $IDD_1*\Delta\, lnSales$+ $IDD_2*\Delta\, lnSales$+ $IDD_r*\Delta\, lnSales$, which is not statistically significant from zero (p=0.532). This result indicates that after the reversal of IDD, the level of cost elasticity for firms with headquarters in states that rejected previous IDD adoptions flips back to the level prior to the initial adoption.

4.2 Robustness Checks

Before we investigate the cross-sectional differences in the post-IDD change in cost elasticity, we conduct several supplemental analyses to test the robustness of our main finding. First, we expand the sample to include IDD adoptions prior to 1977. The data from early years are excluded for our primary analyses because of the concern about the incompleteness of Compustat data as suggested in prior literature (Klasa et al. 2017). Based on the history of IDD adoptions, several states adopted IDD prior to 1977. In addition to New York, which adopted IDD in 1919, the earliest adoption of IDD is by Florida in 1960. We are able to expand the sample to include all firm-year observations starting 5 years before Florida adopted IDD in 1960, i.e. 1955-2011. The results including earlier years (1955-1976) are presented in column (1) of table 5 panel A. The coefficient on $IDD*\Delta\, lnSales$ remains positive and significant at the 1%

level. The economic significance of the IDD impact is also similar to before. The adoption of IDD increases the cost elasticity from 92% to 95% when we expand the sample to include data from early periods.

Second, we limit our sample to the manufacturing industries to examine the robustness of the results. Trade secrets are a very important form of intellectual property for manufacturing companies.[13] We use the SIC code 2000-3999 to identify manufacturing companies and present the results in column (2) of table 5 panel A. Despite the smaller sample size, the coefficient on *IDD*$*\Delta$ *lnSales* remains positive and significant at the 1% level. This result suggests that the IDD adoption also increases the cost elasticity for manufacturing firms. We also find qualitatively similar results when we use the North American industry classification code (NAICS) and classify a firm as in the manufacturing industry if it is between 31 and 33 (untabulated).

Next, we examine whether the variation in certain firm characteristics could explain our main finding. Prior research documents that firms with different levels of employees or assets relative to sales revenues have different levels of cost stickiness (Anderson et al. 2003; Chen et al. 2012). We thus include these two variables to examine whether the impact of IDD on cost elasticity could be driven by changes in asset intensity and/or employee intensity. We also follow Li et al. (2017) to include *Return*, measured as the accumulated stock return over the previous fiscal year, as another control. Higher returns could be a signal of better economic outlook for the firm, which makes it more likely for the firm to keep idle resources, and vice versa. Results after including these variables and their interactions with Δ *lnSales* are presented in Column (3). The results show that firms with higher employee intensity and higher stock returns tend to have higher cost elasticity, and that those with higher asset intensity tend to have

[13] https://www.manufacturinglawblog.com/2016/04/how-passage-of-trade-secrets-act-will-help-manufacturers/

lower cost elasticity. The coefficient on *IDD*Δ lnSales* now captures the incremental effect of IDD adoptions on cost elasticity, after controlling for the influence of employee intensity, asset intensity, and stock returns. The coefficient on *IDD*Δ lnSales* again remains positive and significant at the 1% level.

Lastly, we examine whether other state laws with the potential to change cost behavior during the sample period explain our results. Multiple antitakeover laws are passed in a majority of states starting in 1982. To the extent that these provisions reduce the threat of hostile takeover and empire building, their enactments could affect cost behavior and confound our primary finding (Bertrand and Mullainathan 2003; Chen et al. 2012; Li et al. 2017). We collect the data on the passage of five common types of antitakeover laws from table 2 in Karpoff and Wittery (2017): (1) control share acquisition laws, (2) business combination laws, (3) fair price laws, (4) directors' duties laws, and (5) poison pill laws. Because many states passed several of these in the same year, it may be difficult to test the impact of the incremental effect of each provision. As such, we examine each antitakeover law separately, with the understanding that the effect could in part be attributed to other concurrent laws. We control for the passage of these antitakeover laws using an indicator variable, *Antitakeover*, that is equal to one for firm years when a specific antitakeover law is adopted, and zero otherwise. The indicator variable is interacted with *Δ lnSales.* The results are reported in table 5 panel B. The addition of these antitakeover provisions in the model does not change our main result—the coefficients on *IDD*Δ lnSales* remain positive and significant at the 1% level. The impact of the antitakeover provisions on cost elasticity is positive for all provisions at the 1% level, except for business combination (*BC*), which has a p-value of 0.13. In Column (6), we modify the definition of *Antitakeover* to capture the adoption of any of the five laws and obtain similar results.

4.3 Cross-sectional Variation of the Impact of IDD

In this section, we investigate the cross-sectional variation in the IDD impact on cost elasticity. We define *C* as a variable capturing cross-sectional variation in firm or industry characteristics, and include the interaction of *C* and *Δ lnSales*, and the three-way interaction of *C*, *Δ lnSales*, and *IDD* in the following model.

$$\Delta\, lnCOGS = \beta_0 + \beta_1\, \Delta\, lnSales + \beta_2\, IDD + \beta_3\, IDD{*}\Delta\, lnSales + \beta_4\, GDP\ growth$$
$$+ \beta_5\, GDP\ growth{*}\Delta\, lnSales + \beta_6\, C + \beta_7\, C{*}IDD + \beta_8\, C{*}\Delta\, lnSales + \beta_9 C{*}IDD{*}\Delta\, lnSales$$
$$+ FE + \varepsilon. \qquad (3)$$

The variable of interest in model (3) is *C*IDD*Δ lnSales*, which captures the differential change in cost elasticity following the adoption of IDD, where *C* is one of the four proxies of cross-sectional variation.

Our first cross-sectional partition is based on ex ante demand uncertainty and financial risk faced by the firm. We estimate the standard deviation of sales revenue for each industry defined at the three-digit SIC codes in the previous year, and use it as the proxy for demand uncertainty (*Demand uncertainty*), with a requirement of at least three observations for the calculation. We measure firms' financial strength using z-score (Altman 1968). The indicator variable, *High z-score*, is equal to one if a firm's z-score is above the sample median in the previous year, and zero otherwise. Based on the argument in Section 2, the coefficient on *C *IDD* Δ lnSales* is expected to be positive if *C* represents *Demand uncertainty*, and negative if *C* represents *High z-score*.

The results are presented in table 6. Columns (1) and (2) show the results using *Demand Uncertainty* and *High z-score*, respectively. We first note that the coefficient on *Demand*

*Uncertainty*Δ lnSales* is negative and significant. These results are consistent with the general conclusion from the cost management literature that firms with higher demand uncertainty are more likely to preserve fixed resource capacity to meet future uncertainty, i.e. to main higher operating leverage (Banker et al. 2014a). With regard to the variable of interests, the coefficient on *Demand Uncertainty*IDD*Δ lnSales* is positive and significant at the 1% level, suggesting a higher increase in cost elasticity following the adoption of IDD for firms with higher ex ante demand uncertainty. We also find that the coefficient on *High z-score*IDD*Δ lnSales* is negative and significant at the 5% level, suggesting a more (less) significant increase in cost elasticity following the adoption of IDD for firms with weak (strong) financial strength.

Next, we examine the extent to which the positive impact of IDD on cost elasticity varies with competitive risk. We use two proxies for the competitive risk. The first is measured as the geographical distance between a firm's headquarter and that of all rivals in the industry weighted by each firm's sales (Klasa et al. 2017). The indicator variable, *Far from rivals*, equals to one if the median distance for firms within an industry is above the median in the previous year, and zero otherwise. Given that *Far from rivals* represents lower risk of losing employees with knowledge of trade secrets to rivals, we expect the coefficient on *Far from rivals*IDD*Δ lnSales* to be negative. The other proxy of the competitive risk is associated with the industry entry barriers and product differentiation. High spending in R&D and advertising is assumed to help firms better differentiate their products and provide higher industry entry barriers for potential competitors (Shaked and Sutton 1987; Hoberg and Phillips 2016). We follow Klasa et al. (2017) and use R&D expense plus advertising expense in the previous year deflated by assets as our empirical measure. We use the median value for all firms in a three-digit SIC code in a year as a cutoff point. The indicator variable, *Industry R&D and Advertising*, equals one if a firm is in an industry with higher spending in R&D and advertising than the median, and zero otherwise. Given that *Industry R&D and Advertising* captures lower competitive risk, the coefficient on *Industry R&D and Advertising*IDD*Δ lnSales* is predicted to be negative.

Table 6 column (3) presents the results using the distance to rivals in the industry as a proxy for competitive risk. We note that the coefficient on *Far from rivals* Δ lnSales* is positive and significant, suggesting that firms with lower competitive risk tend to have higher cost elasticity. The coefficient on *Far from rivals* IDD*Δ lnSales* is negative and significant at the 10% level. This supports our prediction that the increase in cost elasticity is weaker for firms in industries with relatively low ex ante risk of losing key employees to rivals.

Table 6 column (4) displays the results using the indicator variable *Industry R&D and Advertising* as another proxy for competitive risk. We again note that the coefficient on *Industry R&D and Advertising * Δ lnSales* is positive and significant, supporting the relationship between cost elasticity and competitive risk documented in the literature. The coefficient on *Industry R&D and Advertising* IDD*Δ lnSales* is negative and statistically significant at the 1%. This is again consistent with our prediction, suggesting that the increase in cost elasticity is smaller for firms in industries with a higher entry barrier and more differentiated products.

4.5 Additional Analysis: Enforcement of Noncompetition Agreements

Finally, we conduct an additional analysis to investigate the interaction of IDD and non-competition agreements, another institutional arrangement that restricts outside employment opportunities for employees. Prior to the adoption of IDD, noncompetition agreements between a firm and its employees are a common form of restrictions on how soon they could be employed by another rival company, usually after one to three years (Bishara, Martin and Thomas 2015;

Schwab and Thomas 2006). There exists considerable variation in the enforceability of such agreements across states, and noncompetition agreements of firms headquartered in states with higher level of enforceability have the potential to provide some protection of trade secrets. To the extent that these states are better able to protect trade secrets, the incremental benefit from the adoption of IDD could be smaller relative to other states with lower enforceability index. We use the enforceability index constructed by Garmaise (2011) based on Malsberger (2004) to examine this possibility.

Because the index is only available between 1992 and 2004, we focus on a subsample between 1993 and 2005 and include the variable, *Enforcement of noncompetition agreements*, measured using the lagged index. The variable is equal to one if the index is above the sample median in the year, and zero otherwise. As shown in Table 7, the coefficient on *Enforcement of noncompetition agreements* IDD*Δ lnSales* is negative and significant at the 5% level, suggesting a muted reaction to IDD for a firm headquartered in a state with higher enforceability of noncompetition. The coefficient on *IDD*Δ lnSales* continues to be positive and significant at the 1% level, but the sum of the two coefficients is not significantly different from zero. This suggests that the increase in cost elasticity as a result of IDD adoptions is concentrated in states with weaker enforcement of noncompetition agreements.

5. Conclusion

In this paper we study the impact of the risk of rival predation arising from a firm's inability to protect trade secrets on cost elasticity. Trade secrets are valuable information existing in all industry sectors, and the inability of firms to protect trade secrets from rivals could expose them to significant risk of predation, and negatively affect profitability and survival.

We build on prior work that highlights the strategic value of preserving fixed resource capacity to meet future uncertainty and to minimize adjusting costs. We posit that greater protection of trade secrets against rivals makes it less likely for rival firms to steal customers and gain market share, thus lowering demand uncertainty, financial risk, and competitive risk. Accordingly, we predict that firms are more likely to reduce the capacity of fixed resources and adopt a more elastic cost structure when there is greater protection of trade secrets.

The empirical setting we use to test the hypothesis is the staggered adoption of the Inevitable Disclosure Doctrine by US state courts over the 1977 to 2011 period. The IDD adoption introduces an arguably exogenous variation that increases the protection of trade secrets, which allows us to conduct a difference-in-differences analysis to study the impact of trade secrets protection on a firm's cost structure.

The key finding is that the recognition of IDD is associated with an increase in cost elasticity in firms headquartered in IDD recognition states relative to those in non-affected states. Further, we find that these results are stronger for firms with higher demand uncertainty and greater financial risk, firms facing more competitive threats due to lower entry barriers and less differentiated products, and firms facing a higher ex-ante risk of losing employees to rival firms. Taken together, these results highlight the strategic value of maintaining fixed resources for firms that face greater risk of losing trade secrets to rival firms.

Our study contributes to an understanding of cost behavior and the determinants of cost elasticity. Our evidence suggests that competitive market conditions have a significant impact on a firm's cost management. A central argument in cost management is that firms respond to exogenous shocks in the environment. The empirical setting used in this paper, the adoption of IDD, provides an exogenous shock that increases the protection of trade secrets and reduces the competitive risk in the market. Therefore, our paper presents new evidence on managers' reactions to outside environmental shocks, which is unlikely a mechanical result of cost behavior.

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Figure 1. The timeline of adopting IDD

This figure is reproduced based on Table 1 of Klasa et al. (2015), which summarizes the calendar year of the precedent-setting legal cases when courts adopted the Inevitable Disclosure Doctrine (IDD). For three states that rejected IDD after initial adopting it (FL in 2001, MI in 2002, and TX in 2003), only the initial adoption years in the graph are shown. The states that are not color-filled in the figure either never considered or considered but rejected the IDD.

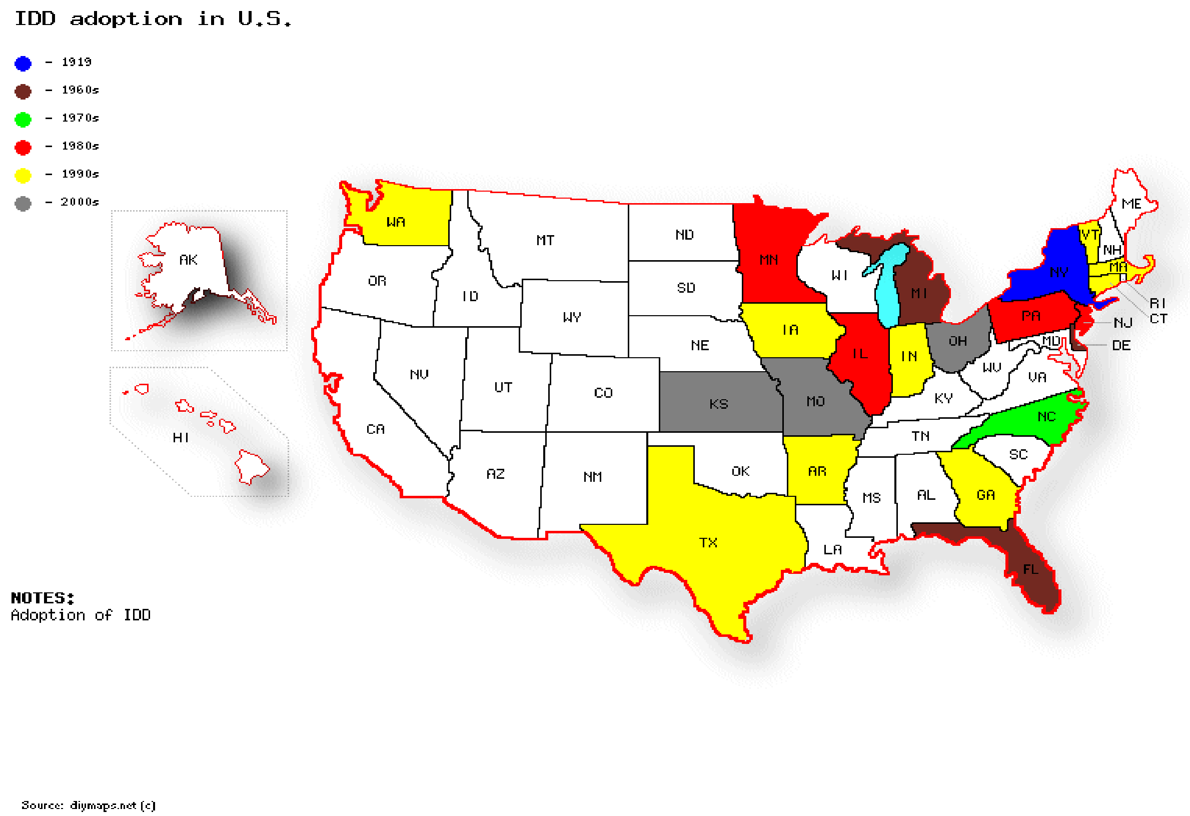

Table 1. Sample selection

This table shows how we get the final sample for our main analysis starting from the merged file of Compustat-CRSP.

	Firm Year obs
All Compustat-CRSP merged U.S. Firms from 1977 to 2011	139,853
Less: regulated industries like utilities and financial firms	116,884
Less: missing prior period COGS or Sales	106,345
Less: extreme values, 1% of sales or COGS at both tails for each year	102,407

Table 2. Summary Statistics

This table present summary statistics for the sample used in all models. Δ *lnCOGS* is the change in the natural logarithm of cost of goods sold for a firm from year *t*-1 to year *t*. Δ *lnSales* is the change in the natural logarithm of sales revenue for a firm from year *t*-1 to year *t*. *IDD* is an indicator variable equal to one if a firm's headquarter is in a state that has adopted IDD as of year *t*, and zero otherwise. *GDP growth* is the percentage change in U.S. GDP measured in 2009 dollars from year *t*-1 to year *t*. *Employee intensity* is the ratio of number of employees to sales revenue in year *t*-1. *Asset intensity* is the ratio of total assets to sales revenue in year *t*-1. *Return* is the 12-month buy-and-hold return for the firm in the fiscal year *t*-1, ending in three months after the fiscal year end. *Demand Uncertainty* is the standard deviation of sales revenue in a firm's industry (based on the 3-digit SIC code) in year *t*-1, with a minimum of three firms in each industry. *High z-score* is an indicator variable equal to one if a firm's z-score in year *t*-1 is above sample median, and zero otherwise. *Far from rivals* is an indicator variable equal to one if the median distance to all rival firms in an industry, weighted by sales revenue for each rival firm, is above the sample median in year *t*-1, and zero otherwise. *Industry R&D and Advertising* is an indicator variable equal to one if the median ratio of R&D and advertising in a firm's industry has above the median spending in R&D and Advertising deflated by total assets in year *t*-1 and zero otherwise. *Noncompetition index* is an indicator variable equal to one if the index for the enforcement of noncompetition agreements in the stats a firm is headquartered in for year *t*-1 is above the sample median, and zero otherwise. The index is compiled by Garmaise (2011) and available between 1992 and 2012.

Variable	N	Mean	Median	Std dev
Δ *lnCOGS*	102,407	0.101	0.089	0.225
Δ *lnSales*	102,407	0.097	0.088	0.214
IDD	102,407	0.417	0.000	0.493
GDP growth	102,407	0.030	0.035	0.020
Employee intensity	99,789	0.011	0.008	0.011
Asset intensity	99,789	1.032	0.796	0.822
Return	99,789	0.170	0.061	0.612
Demand Uncertainty	98,842	0.195	0.193	0.072
High z-score	94,366	0.500	0.000	0.500
Far from rivals	102,407	0.358	0.000	0.479
Industry R&D and Advertising	102,407	0.794	1.000	0.404
Noncompetition index	41,866	0.400	0.000	0.490

Table 3. The impact of IDD adoption on cost elasticity

This table reports the regression results of the impact of IDD adoption on cost elasticity based on the sample 1977-2011. Δ *lnCOGS* is the change in the natural logarithm of cost of goods sold for a firm from year *t*-1 to year *t*. Δ *lnSales* is the change in the natural logarithm of sales revenue for a firm from year *t*-1 to year *t*. *IDD* is an indicator variable equal to one if a firm's headquarter is in a state that has adopted IDD as of year *t*, and zero otherwise. *GDP growth* is the percentage change in U.S. GDP measured in 2009 dollars from year *t*-1 to year *t*. ***, **, * denote statistical significance at the 1%, 5% and 10% levels for two-sided tests. Standard errors are clustered by firm.

Dependent variable = Δ *lnCOGS*			
	Coefficient (std err) (1)	Coefficient (std err) (2)	Coefficient (std err) (3)
Δ *lnSales*	0.910***	0.910***	0.908***
	(0.005)	(0.005)	(0.006)
IDD	-0.005***	-0.003***	-0.002
	(0.001)	(0.001)	(0.001)
IDD *Δ *lnSales*	0.034***	0.034***	0.032***
	(0.005)	(0.005)	(0.006)
GDP growth	-0.048	-0.050	-0.054
	(0.037)	(0.037)	(0.038)
*GDP growth**Δ *lnSales*	0.100	0.106	-0.009
	(0.108)	(0.108)	(0.124)
Fixed effects	Year, Industry	Year, Industry, State	Year, Firm
adj.R2	0.793	0.793	0.825
N	102,407	102,407	102,407

Table 4. The change in COGS around the adoption of IDD

This table reports the regression results of the impact of IDD adoption on cost elasticity based on the sample 1977-2011. Δ *lnCOGS* is the change in the natural logarithm of cost of goods sold for a firm from year *t*-1 to year *t*. Δ *lnSales* is the change in the natural logarithm of sales revenue for a firm from year *t*-1 to year *t*. IDD_0, which equals to one if a firm's headquarter is in a state that adopts the IDD in the current year. IDD_{-1} and IDD_{-1} are indicator variables equal to one if a firm's headquarter is in a state that adopted the one year ago and two year ago, and zero otherwise. IDD_1 and IDD_2 are equal to one if a firm's headquarter is in a state that will adopt the IDD in one year and in two years. IDD_r equals to one starting from the first year of rejection, and zero otherwise. *GDP growth* is the percentage change in U.S. GDP measured in 2009 dollars from year *t*-1 to year *t*. ***, **, * denote statistical significance at the 1%, 5% and 10% levels for two-sided tests. Standard errors are clustered by firm.

	Coefficient (std err) (1)	Coefficient (std err) (2)	Coefficient (std err) (3)
Δ *lnSales*	0.926***	0.923***	0.920***
	(0.004)	(0.004)	(0.005)
IDD_0	-0.001	0.001	-0.001
	(0.003)	(0.003)	(0.003)
IDD_0 *Δ *lnSales*	0.032**	0.032**	0.031*
	(0.014)	(0.014)	(0.018)
IDD_1	-0.000	0.001	0.000
	(0.003)	(0.003)	(0.003)
IDD_1 *Δ *lnSales*	0.022	0.022	0.021
	(0.014)	(0.014)	(0.015)
IDD_2	-0.002	-0.001	-0.002
	(0.003)	(0.003)	(0.003)
IDD_2 *Δ *lnSales*	0.012	0.011	0.018
	(0.014)	(0.014)	(0.015)
IDD_{-1}	0.001	0.002	0.001
	(0.003)	(0.003)	(0.003)
IDD_{-1} *Δ *lnSales*	0.008	0.008	0.006
	(0.014)	(0.014)	(0.017)
IDD_{-2}	-0.000	0.000	-0.001
	(0.003)	(0.003)	(0.003)
IDD_{-2} *Δ *lnSales*	0.020	0.022	0.025
	(0.015)	(0.015)	(0.016)
IDD_r	-0.000	-0.003	-0.006*
	(0.002)	(0.003)	(0.003)
IDD_r *Δ *lnSales*	-0.045***	-0.049***	-0.039**
	(0.016)	(0.016)	(0.019)
GDP growth	-0.045	-0.049	-0.053
	(0.037)	(0.037)	(0.038)
*GDP growth**Δ *lnSales*	0.063	0.089	-0.027
	(0.109)	(0.109)	(0.125)
Fixed Effects	Year, Industry	Year, Industry, State	Year, Firm
adj.R2	0.792	0.793	0.824
N	102,407	102,407	102,407

Table 5. Robustness checks

Panel A. Robustness checks

This table reports the robustness checks for the impact of IDD adoption on cost elasticity. Column (1) expands the sample to 1955-2011. Column (2) focuses on firms in the manufacturing industry. Column (3) adds additional control variables. Δ *lnCOGS* is the change in the natural logarithm of cost of goods sold for a firm from year *t*-1 to year *t*. Δ *lnSales* is the change in the natural logarithm of sales revenue for a firm from year *t*-1 to year *t*. *IDD* is an indicator variable equal to one if a firm's headquarter is in a state that has adopted IDD as of year *t*, and zero otherwise. *GDP growth* is the percentage change in U.S. GDP measured in 2009 dollars from year *t*-1 to year *t*. *Employee intensity* is the ratio of number of employees to sales revenue in year *t*-1. *Asset intensity* is the ratio of total assets to sales revenue in year *t*-1. *Return* is the 12-month buy-and-hold return for the firm in the fiscal year *t*-1, ending in three months after the fiscal year end. ***, **, * denote statistical significance at the 1%, 5% and 10% levels for two-sided tests. Standard errors are clustered by firm.

Dependent variable = Δ *lnCOGS*			
	Coefficient (std err) 1955-2011 (1)	Coefficient (std err) Manufacturing firms (2)	Coefficient (std err) Add controls (3)
Δ *lnSales*	0.915***	0.914***	0.992***
	(0.005)	(0.008)	(0.007)
IDD	-0.001	-0.000	-0.000
	(0.001)	(0.001)	(0.001)
IDD *Δ *lnSales*	0.024***	0.021***	0.019***
	(0.005)	(0.007)	(0.006)
GDP growth	-0.062**	-0.100	-0.056
	(0.029)	(0.083)	(0.037)
*GDP growth**Δ *lnSales*	0.062	0.411**	-0.066
	(0.104)	(0.195)	(0.126)
Employee intensity			0.003
			(0.107)
*Employee intensity**Δ *lnSales*			0.921***
			(0.236)
Asset intensity			0.027***
			(0.002)
*Asset intensity**Δ *lnSales*			-0.076***
			(0.005)
Return			-0.003***
			(0.001)
*Return**Δ *lnSales*			0.014***
			(0.004)
Fixed effects	Year, Firm	Year, Firm	Year, Firm
adj.R2	0.830	0.867	0.831
N	125,682	41,774	99,789

Panel B. Confounding antitakeover provisions

This table examines the robustness of the impact of IDD adoption on cost elasticity after controlling for antitakeover provisions. The first five columns control for the adoption of each of the five antitakeover provisions. For example, in Column (1), the variable *Antitakeover*, is equal to one if a firm's headquarter state adopts control shares acquisition laws (CS) as of year *t*, and zero otherwise. Column (6) combines the effect of all five provisions. Δ *lnCOGS* is the change in the natural logarithm of cost of goods sold for a firm from year *t*-1 to year *t*. Δ *lnSales* is the change in the natural logarithm of sales revenue for a firm from year *t*-1 to year *t*. *IDD* is an indicator variable equal to one if a firm's headquarter is in a state that has adopted IDD as of year *t*, and zero otherwise. *GDP growth* is the percentage change in U.S. GDP measured in 2009 dollars from year *t*-1 to year *t*. *CS* is an indicator variable equal to one for the adoption of control share acquisition laws; *BC* is an indicator variable equal to one for the adoption of business combination laws; *FP* is an indicator variable equal to one for the adoption of fair price laws; *DD* is an indicator variable equal to one for the adoption of directors' duties laws; *PP* is an indicator variable equal to one for the adoption of poison pill laws. ***, **, * denote statistical significance at the 1%, 5% and 10% levels for two-sided tests. Standard errors are clustered by firm.

Antitakeover =	*CS*	*BC*	*FP*	*DD*	*PP*	*Any provision*
	Coefficient	Coefficient	Coefficient	Coefficient	Coefficient	Coefficient
	(std err)	(std err)	(std err)	(std err)	(std err)	(std err)
	(1)	(2)	(3)	(4)	(5)	(6)
Δ *lnSales*	0.904***	0.905***	0.897***	0.901***	0.904***	0.904***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
IDD	-0.002*	-0.002	-0.000	-0.001	-0.002	-0.002
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
IDD *Δ *lnSales*	0.030***	0.027***	0.011*	0.025***	0.028***	0.028***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
GDP growth	-0.055	-0.055	-0.053	-0.057	-0.056	-0.055
	(0.038)	(0.038)	(0.038)	(0.038)	(0.038)	(0.038)
*GDP growth**Δ *lnSales*	0.002	0.008	0.019	0.036	0.025	0.000
	(0.125)	(0.124)	(0.124)	(0.124)	(0.124)	(0.124)
Antitakeover	-0.005***	0.002	-0.001	-0.003**	-0.001	0.000
	(0.001)	(0.002)	(0.002)	(0.001)	(0.001)	(0.002)
*Antitakeover**Δ *lnSales*	0.017***	0.010	0.049***	0.022***	0.012**	0.009
	(0.007)	(0.006)	(0.006)	(0.006)	(0.006)	(0.007)
Fixed effects	Year, Firm	Year, Firm	Year, Firm	Year, Firm	Year, Firm	Year, Firm
adj.R2	0.825	0.825	0.825	0.825	0.825	0.825
N	102,407	102,407	102,407	102,407	102,407	102,407

Table 6. Cross-sectional variations of the impact of IDD adoption

This table reports results estimating the cross-sectional variations of the impact of IDD adoption on cost elasticity. Δ *lnCOGS* is the change in the natural logarithm of cost of goods sold for a firm from year *t*-1 to year *t*. Δ *lnSales* is the change in the natural logarithm of sales revenue for a firm from year *t*-1 to year *t*. *IDD* is an indicator variable equal to one if a firm's headquarter is in a state that has adopted IDD as of year *t*, and zero otherwise. *GDP growth* is the percentage change in U.S. GDP measured in 2009 dollars from year *t*-1 to year *t*. *Demand Uncertainty* is the standard deviation of sales revenue in a firm's industry (based on the 3-digit SIC code) in year *t*-1, with a minimum of three firms in each industry. *High z-score* is an indicator variable equal to one if a firm's z-score in year *t*-1 is above sample median, and zero otherwise. *Far from rivals* is an indicator variable equal to one if the median distance to all rival firms in an industry, weighted by sales revenue for each rival firm, is above the sample median in year *t*-1, and zero otherwise. *Industry R&D and Advertising* is an indicator variable equal to one if the median ratio of R&D and advertising in a firm's industry has above the median spending in R&D and Advertising deflated by total assets in year *t*-1 and zero otherwise. ***, **, * denote statistical significance at the 1%, 5% and 10% levels for two-sided tests. Standard errors are clustered by firm.

Dependent variable = Δ *lnCOGS*				
C =	*Demand Uncertainty*	*High Z-score*	*Far from rivals*	*Industry R&D and Advertising*
	Coefficient (std err) (1)	Coefficient (std err) (2)	Coefficient (std err) (3)	Coefficient (std err) (4)
Δ *lnSales*	1.004***	0.906***	0.903***	0.875***
	(0.011)	(0.007)	(0.007)	(0.010)
IDD	0.003	-0.001	-0.003**	-0.005***
	(0.002)	(0.002)	(0.001)	(0.002)
*IDD*Δ lnSales*	-0.011	0.042***	0.040***	0.066***
	(0.015)	(0.009)	(0.008)	(0.011)
C	0.036***	0.020***	-0.001	-0.001
	(0.010)	(0.002)	(0.002)	(0.002)
*C*IDD*	-0.021*	-0.001	0.003	0.000
	(0.013)	(0.002)	(0.002)	(0.002)
*C*Δ lnSales*	-0.464***	0.010	0.015*	0.046***
	(0.053)	(0.008)	(0.009)	(0.010)
*C*IDD*Δ lnSales*	0.188**	-0.021**	-0.020*	-0.040***
	(0.074)	(0.011)	(0.011)	(0.012)
GDP growth	-0.092**	-0.052	-0.053	-0.041
	(0.038)	(0.037)	(0.038)	(0.037)
*GDP growth*Δ lnSales*	0.208*	-0.076	-0.003	0.045
	(0.125)	(0.127)	(0.124)	(0.108)
Fixed effects	Year, Firm	Year, Firm	Year, Firm	Year, Firm
adj.R2	0.826	0.835	0.825	0.793
N	98,842	94,366	102,407	102,407

Table 7. IDD and enforceability of noncompetition agreements

This table reports whether the impact of IDD adoption on cost elasticity depends on variations in the enforceability of noncompetition agreements.Δ *lnCOGS* is the change in the natural logarithm of cost of goods sold for a firm from year *t*-1 to year *t*. Δ *lnSales* is the change in the natural logarithm of sales revenue for a firm from year *t*-1 to year *t*. *IDD* is an indicator variable equal to one if a firm's headquarter is in a state that has adopted IDD as of year *t*, and zero otherwise. *GDP growth* is the percentage change in U.S. GDP measured in 2009 dollars from year *t*-1 to year *t*. *Noncompetition index* is an indicator variable equal to one if the index for the enforcement of noncompetition agreements in the stats a firm is headquartered in for year *t*-1 is above the sample median, and zero otherwise. It is compiled by Garmaise (2011) and available between 1992 and 2012. ***, **, * denote statistical significance at the 1%, 5% and 10% levels for two-sided tests. Standard errors are clustered by firm.

Dependent variable = Δ *lnCOGS*	
	Coefficient (std err)
Δ *lnSales*	0.891*** (0.015)
IDD	0.006 (0.004)
*IDD**Δ *lnSales*	0.029** (0.012)
Noncompetition Index	0.001 (0.004)
*Noncompetition Index*IDD*	-0.004 (0.004)
*Noncompetition Index**Δ *lnSales*	0.044*** (0.014)
*Noncompetition Index*IDD**Δ *lnSales*	-0.036** (0.018)
GDP growth	-0.361*** (0.115)
*GDP growth**Δ *lnSales*	0.350 (0.351)
Fixed effects	Year, Firm
adj.R2	0.840
N	41,864