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Do labor mobility restrictions affect debt maturity?

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ABSTRACT

Prior literature finds that staggered state-level adoption of the Inevitable Disclosure Doctrine (IDD) significantly constrains labor mobility. Using the IDD as an exogenous shock to labor mobility, we find that firms headquartered in states that adopt the IDD gravitate towards issuing short-term debt for external debt financing. We examine three mechanisms—default risk, information asymmetry, and agency cost mitigation—through which labor mobility restrictions affect debt maturity. Our results provide support for the information asymmetry mechanism, which suggests that firms are more inclined to use short-term debt when their information environment deteriorates. We find that in the wake of IDD adoption, firms tend to utilize short-term debt only in corporate bond markets and their debt maturity profiles become more concentrated.

1. Introduction

The diminishing value of preserving unused debt capacity leads firms to leverage up in the face of increased labor mobility restrictions (Klasa et al., 2018). As the threats of leaking valuable trade secrets to rivals increase, firms hold unused debt capacity to defend their competitive position and signal their financial flexibility to their rivals if attacked (Klasa et al., 2018). However, labor mobility restrictions help firms protect their trade secrets more effectively, incentivizing them to use more debt since holding unused debt capacity becomes less necessary as competitive threats decrease.

Increased leverage is known to aggravate shareholder-creditor conflicts (Myers, 1977), which firms can mitigate using short-term debt to alleviate such suboptimal effects as underinvestment and asset substitution problems (Myers, 1977; Barnea et al., 1980; Leland and Toft, 1996). Short-term debt, however, can incur liquidity risk due to a possible maturity mismatch between sources and uses of funds (Diamond, 1991), thereby exposing firms to potential inefficient liquidation of positive NPV projects and lowering their value. This begs the question of whether an increase in leverage assumes the form of short-term or long-term debt.

Using U.S. state courts' staggered adoption of the IDD as an exogenous shock to labor mobility, this paper investigates the effects of labor mobility restrictions on firms' debt maturity choice, and finds firms headquartered in states that adopt the IDD to gravitate towards issuing short-term debt for external financing. We extend the findings in Klasa et al. (2018) by focusing on the relation between labor mobility restrictions and choice of debt maturity, the latter being an important feature of debt that affects both investment and liquidity policy in firms that increase their debt level.

We propose and examine two competing views on the association between labor mobility restrictions and debt maturity. The first maintains that labor mobility restrictions that limit outside employment opportunities, because they heighten employees' career concerns, render managers more risk averse (Islam et al., 2019) and long-term oriented (Islam et al., 2021). Risk-averse managers may prefer long-term debt to avoid the frequent scrutiny and inefficient liquidation that can attend the need to refinance debt. Firms focused on long-term performance are more likely to employ long-term debt to fund long-term investments (Guedes and Opler, 1996).

The competing view proposes three reasons why labor mobility restrictions can shorten debt maturity. The first is that a focus on long-term performance and greater managerial risk aversion consequent to a reduction in labor mobility reduce the risk of default. Firms with lower default risk may prefer short-term debt because of low interest rates and low liquidation risk (Diamond, 1991). Second, limiting labor mobility is associated with increased information asymmetry because the legal remedies available under the IDD help prevent leakage of proprietary information to competitors and thereby increase the cost of proprietary information disclosure (Aobdia, 2018; Li et al., 2018; Callen et al.,

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2020). Short-term debt is preferred because it exposes managers to greater scrutiny by creditors, which reduces the effect of private information on the cost of financing (Guedes and Opler, 1996). Further, being less sensitive to underpricing, short-term debt may be used by firms with underpriced liabilities to signal to investors their true value (Guedes and Opler, 1996). Third, because labor mobility restrictions reduce product market competitive threats by preventing leakage of firms' proprietary information to rivals (Klasa et al., 2018), short-term debt may substitute for less external monitoring by the product market.

Taken together, the trade-offs of labor mobility restrictions and benefits and costs associated with choosing short-term over long-term debt render the net effect of labor mobility restrictions on debt maturity ambiguous. Given the opposing effects on debt maturity elucidated above, the relationship between labor mobility restrictions and debt maturity remains an empirical question. We use a difference-in-differences (DiD) research design to analyze a sample of 75,126 firm-year observations for U.S. publicly listed firms from 1977 to 2015. Debt maturity is measured as the proportion of long-term debt maturing more than five years. Due to the staggered adoption of IDD, our treatment group is composed of firms headquartered in adopting, the control group of firms headquartered in non-adopting, states. Our finding of a significant decrease in long-term debt financing following adoption of the IDD confirms a positive relation between labor mobility restrictions and short-term debt financing. Our results are robust to using alternative measures of debt maturity.

A battery of tests performed to confirm the robustness of our main results includes a placebo test with fictitious IDD changes and dynamic analysis to allay the concern that our results might be driven by non-parallel trends and simultaneity, and propensity score matching to address concerns related to selection bias. We re-run the baseline regression using a stacked regression to address potential biases associated with the staggered DiD setup, re-estimate the baseline model including two state-level legal variables, and perform a neighboring-state test to mitigate concerns related to state-level confounding factors. The results of these tests confirm the robustness of our findings.

We next investigate whether the positive effect of labor mobility restrictions on short-term debt financing can be explained by the default risk, information asymmetry, and agency cost mitigation mechanisms. We find no support for the default risk mechanism, using as proxies for default risk Altman's (1968) z-score and the KZ index (Kaplan and Zingales, 1997). Examining the information asymmetry mechanism using analyst forecast dispersion and analyst forecast error as proxies, we find that firms tend to increase short-term debt financing following adoption of the IDD due to the deteriorating information environment. Examining the non-monotonic relationship between default risk and debt maturity by splitting the full sample into extreme and moderate default risk groups using the z-score and KZ index separately, we find that the interaction terms between the proxy for information asymmetry and IDD indicator variable are statistically significant for the extreme (low and high) default risk subsample. This result provides evidence that supports Diamond's (1991) finding that low and high (moderate) default risk firms tend to use short-term (long-term) debt.

With respect to the agency cost mitigation mechanism, we find no evidence that firms substitute short-term debt for reduced external monitoring stemming from product market competitive threats. Overall, our results suggest that firms hold more short-term debt after IDD adoption because doing so can help to mitigate information asymmetry between shareholders and creditors, which supports the information asymmetry mechanism.

Additional analyses reveal the effect of IDD adoption on short-term debt financing to be magnified in high innovation-intensive firms and firms that face geographically proximate rivals that can more easily poach their workers. We further find the effect of IDD adoption on debt maturity to be observed only in firms that face financially stronger rivals and heightened competitive threats due to higher asset specificity.

Because the balance sheet approach to measuring debt maturity is based on historical aggregations and reflects past managerial decisions about firms' capital structure, we use new debt issue (that is, the incremental approach) to examine the impact of labor mobility restrictions across all points along the spectrum of debt maturity (Guedes and Opler, 1996; Custodio et al., 2013). This approach facilitates investigation of how the debt maturity structure changes following adoption of the IDD from the perspective of a prospective creditor. Our analysis, which utilizes both new bond issues from the Mergent Fixed Income Securities Database (FISD) and syndicated loan data from the Dealscan database, reveals firms to tend to shift away from long-term debt after IDD adoption in corporate bond markets, corroborating our baseline findings. No such effect is observed in private debt markets.

Having established a negative relation between labor mobility restrictions and debt maturity, we further examine whether increased labor mobility restrictions result in a more concentrated debt maturity profile. An optimal debt maturity profile is determined by the trade-off between increasing rollover risk arising from concentrated debt maturity and the low cost of a few large debt issues (Choi et al., 2018). Our finding that firms' debt maturity dispersion is reduced following IDD adoption implies that restricting labor mobility is associated with a more concentrated debt maturity profile. Increased labor mobility restrictions evoke a stronger response in high-leverage firms, which reduce their debt maturity dispersion, than in low-leverage firms.

Our study contributes to the existing literature in two ways. We contribute to the burgeoning literature on debt structure the introduction of a new factor that can affect debt maturity, namely, restrictions on labor mobility consequent to adoption of the IDD. Prior empirical studies show corporate governance (Harford et al., 2008), growth opportunities (Barclay and Smith, 1995; Guedes and Opler, 1996), executive compensation (Brockman et al., 2010; Dang and Phan, 2016), managerial ownership (Datta et al., 2005), CEO overconfidence/managerial ability (Huang et al., 2016; Shang, 2021), threat of entry (Parise, 2018), asset salability (Benmelech, 2008), policy uncertainty (Datta et al., 2019), financial crisis (Gonzalez, 2015), national culture (Zheng et al., 2012), and macroeconomic conditions (Erel et al., 2012) to affect firms' debt maturity decisions. Our analysis documents firms to gravitate towards short-term debt financing following an increase in labor mobility restrictions resulting from state recognition of the IDD. Using short-term debt helps to mitigate the information asymmetry generated by limiting labor mobility.

We also extend the literature on the economic and labor market effects of restricting labor mobility via trade secret protection laws, in particular, adoption of the IDD (e.g., Klasa et al., 2018; Li et al., 2018; Qiu and Wang, 2018; Na, 2021), by showing that limiting labor mobility also affects firms' debt maturity through increased information asymmetry.

The remainder of the paper is organized as follows. Section 2 describes the IDD, its relation to labor mobility, and the theoretical link between labor mobility restrictions and short-term debt financing. Section 3 describes our data sources and variables construction. Section 4 presents the empirical research design and results of the baseline models and robustness tests. Section 5 examines the mechanisms through which labor mobility restrictions affect debt maturity; Section 6 discusses additional tests. We report the results of a new debt maturity analysis in Section 7 and in Section 8 describe the analysis of debt maturity dispersion. Section 9 concludes.

2. Background and hypothesis development

In this section, we provide a brief introduction to trade secret protection laws and their relation to labor mobility. We then review the related literature to establish the theoretical link between labor mobility restrictions and short-term debt financing.

2.1. Trade secret protections and labor mobility

A trade secret is information not generally known to competitors that provides economic advantage to its owner and warrants reasonable effort to maintain it as secret (Uniform Trade Secret Act, 1996). Examples include formulas, patterns, compilations of customer information, computer programs, and manufacturing processes. Ownership and protection of a trade secret can yield tangible benefits, such as increased profitability and improved shareholder value, misappropriation exerts a deleterious impact on the owner. Viewed as arising from the notion of "threatened misappropriation" in the Uniform Trade Secret Act (UTSA) (Wiesner, 2012), the IDD permits courts to enjoin departing employees from working for their employers' competitors if it can be demonstrated that "the employee's new job duties will inevitably cause the employee to rely upon knowledge of the former employer's trade secrets" (*Whyte v. Schlage Lock Co.,* 2002).[1] Application of the IDD is based on the threat of misappropriation. In other words, the IDD may be applicable even in the absence of actual misappropriation and a non-compete agreement (Matheson, 1998). *PepsiCo* also sets the standard for evaluating the existence of inevitable disclosure. Not all U.S. states apply the IDD, and of those that do, some have adopted it in its entirety, others with limitations in application.

In restraining employees from working for their employers' competitors or starting competing firms, the IDD limits outside employment opportunities by reducing the information that can be brought to a new employer. Employees subject to the IDD are thus less likely to move to other firms, especially in the same industry. Prior literature provides evidence to support the IDD's effectiveness in reducing labor mobility. For example, the mobility for all workers is lower for states with than for those without the IDD legal precedent (Png and Samila, 2013), and the employees in managerial and related positions are less likely to move to rival firms in the same state as well as to firms in other states that adopt the IDD (Klasa et al., 2018).

In this paper, we employ the staggered state adoption of the IDD as the plausible exogenous shock to labor mobility. We use the precedent-setting cases adopting and/or rejecting the IDD compiled by Qiu and Wang (2018). The state adoption of the IDD is suitable as the quasi-natural experiment for our debt maturity test for the following three reasons. First, the judicial decision for the application of this doctrine in the precedent-setting cases is not aimed to affect the firm's debt maturity. This is because in applying the IDD, the state court must strike a balance between protecting employers' trade secret and preserving the employee's freedom of employment right (Harris, 2000; Godfrey, 2004).

Second, unlike the passage or defeat of the state or federal law that could be influenced by the lobbying interest groups, the adoption or rejection of the doctrine has been judicial decisions that are driven mainly by the merits of trade secret misappropriation claim in each legal case (Klasa et al., 2018). Accordingly, the state court's decision on the application of the IDD is unlikely to be anticipated by firms. Indeed, Klasa et al. (2018) find that apart from the covenant not to compete, all other factors including the state-level factors related to labor and trade secrets laws such as the passage of wrongful discharge law, right-to-work laws, UTSA, union membership, characteristics of the workforce (college enrolment and age of workers), local economic and political conditions, and shock to the state's legal and business environment do not affect the likelihood of the IDD adoption. Third, the scope of protection provided by the IDD is much broader than the one by the non-compete agreement and UTSA. By the IDD, an injunction can be obtained on the ground that the firm (trade secret owner) can demonstrate that the departing employee would inevitably disclose the trade secrets. The firm is not required to prove that the departing employee had used or disclosed or threatened to disclose the trade secret (Png and Samila, 2013).

2.2. Labor mobility restrictions and short-term debt financing

The optimal debt maturity structure is determined by balancing the benefits and costs of using short-term debt. Benefits include lower debt mispricing risk occasioned by information asymmetry and lower agency costs of underinvestment and asset substitution. Costs, include increased liquidity risk. In this section, we elucidate competing views of the theoretical link between labor mobility restrictions and short-term debt financing.

2.2.1. Negative relation between labor mobility restrictions and short-term debt financing

Prior studies find a state's adoption of the IDD to affect employee behavior by reducing outside employment opportunities. Labor mobility restrictions render managers more risk averse by constraining outside career opportunities. To safeguard their current positions, less mobile, risk-averse managers tend to prefer more conservative corporate policies that may not be optimal for shareholders (Fama, 1980; Colak and Korkeamaki, 2021), such as long-term over short-term debt, the former limiting exposure to rollover risk and reducing scrutiny by creditors (Dang and Phan, 2016). Firms with riskier growth opportunities, in particular, are incentivized to lengthen debt maturity to avoid inefficient liquidation resulting from failure to roll over short-term debt (Guedes and Opler, 1996). Labor mobility restrictions may also lead managers to become more long-term oriented and make more long-term investments (Islam et al., 2021), the latter more likely to be funded using long-term debt. Taken together, the foregoing arguments yield the following hypothesis.

H1a: Labor mobility restrictions are negatively associated with short-term debt financing.

2.2.2. Positive relation between labor mobility restrictions and short-term debt financing

Restricting labor mobility can increase short-term debt financing via three mechanisms, default risk, information asymmetry, and agency cost mitigation.

Underlying the default risk mechanism are three arguments. First, a less mobile labor market reduces exposure to competitive threats by denying rivals easy access to a firm's trade secrets by means of poaching employees. Diminished competitive threats can translate to lower default risk (Choi, 2020). Second, labor mobility restrictions serve to bind employees to firms, thereby reducing the costs both of hiring and of debt financing (Png and Samila, 2013; Choi, 2020). Third, restricting labor mobility can motivate managers desirous of keeping their current jobs secure to be more conservative (Islam et al., 2019) and focus on long-term performance rather than myopic, short-term goals (Islam et al., 2021). This, in turn, strengthens firms' competitive positions and reduces their default risk. Because low default risk improves access to external debt financing, managers in firms located in states adopting the IDD may be less concerned about rollover risk and incentivized to choose short-term debt at low interest rates. This is consistent with Diamond's (1991) theoretical findings that low-risk firms tend to use short-term debt.

The second mechanism is based on the argument that adoption of the IDD, by restricting labor mobility, increases information asymmetry by reducing the risk of firms losing confidential information to rivals (Aobdia, 2018; Li et al., 2018). Debt may be mispriced to the extent that lenders require a higher yield to invest in debts issued by more informationally opaque firms. Previous studies find short-term debt to help attenuate information asymmetry between shareholders and creditors (Flannery, 1986; Diamond, 1991). Two models—the adverse selection and signalling models (Guedes and Opler, 1996)—have been proposed to explain how short-term debt can be expected to mitigate information asymmetry. According to the adverse selection model, short-term debt,

[1] 101 Cal.App.4th 1447.

because it requires frequent recontracting, exposes managers to greater scrutiny by creditors (Diamond, 1991). Firms with high information asymmetry may be amenable to the additional scrutiny imposed by short-term debt because absence of credible governance to reduce information asymmetry increases the cost of capital when raising external equity.

The signalling model suggests that short-term debt may be preferred by more informationally opaque firms to signal to the markets the true value of good projects (Flannery, 1986). Firms with underpriced liabilities gravitate towards issuing short-term debt because it is less sensitive to underpricing (Guedes and Opler, 1996). Accordingly, if labor mobility restrictions expose firms to higher information asymmetry, we expect firms headquartered in states that adopt the IDD, and that possess favorable information about their default risk, to have more short-term debt.

Labor mobility restrictions reduce debt maturity, the third mechanism, by helping to ameliorate agency costs. Passage of the IDD helps to preserve or strengthen firms' market power by reducing the likelihood of losing business to competitors through information leakage (Klasa et al., 2018). Reduced competitive threat translates to less external monitoring by the product market, as a substitute for which firms may exhibit a greater need for the frequent monitoring provided by short-term debt.

The foregoing arguments suggest the following hypothesis.

H1b: Labor mobility restrictions are positively associated with short-term debt financing.

3. Sample and variables description

Our primary sample includes U.S. public listed firms for the period 1977 – 2015.[2] We begin with the firms in the Compustat database, excluding firms from the financial (SIC 4900–4999) and utility (SIC 6000–6999) industries. We obtain stock return data from the Center for Research in Security Prices (CRSP). Merging Compustat data and CRSP data yields 75,126 firm-year observations with no missing values for key dependent and independent variables. We collect analyst forecast data from Institutional Brokers' Estimation System (I/B/E/S), credit rating data from Compustat, new bond issues data from Mergent FISD, new loans data from Reuter DealScan, and debt maturity dispersion data from Standard & Poor's Capital IQ. We winsorize continuous variables to remove outliers at their 1st and 99th percentiles.

3.1. Dependent variable: debt maturity

The key dependent variable of debt maturity (*LT5*) is defined as the ratio of long-term debt minus debt maturing in 2, 3, 4, and 5 years to total debt (Datta et al., 2005; Chen et al., 2013; Custodio et al., 2013). As a robustness check, we follow Custodio et al. (2013) and Huang et al. (2016) in employing four alternative measures of debt maturity: *LT4* to *LT1*, the proportion of long-term debt maturing more than 4 years to the proportion of long-term debt maturing more than 1 year, respectively.

3.2. Independent variable

Following Klasa et al. (2018), we use a precedent-setting IDD case in a state to construct the independent variable *IDD*, an indicator variable that equals one if a state has adopted the IDD during and after any given year, and zero otherwise.[3] If a court decision reverses a state's initial IDD adoption position, the *IDD* equals zero during and after the rejection year of the IDD court's decision, and one otherwise. For states with case law that have not explicitly considered or rejected the *IDD*, the *IDD* is equal to zero for each year of the sample period. We adopt the list of IDD adoption/rejection court decisions in Qiu and Wang (2018).[4] Table 1 presents the timeline of IDD adoption/rejection court decisions by state that are relevant to our sample period. Of 32 court cases that have adopted or rejected the IDD, 23 are adopting and 9 rejecting court cases.

3.3. Control variables

Following previous studies on debt maturity (e.g., Stohs and Mauer, 1996; Johnson, 2003; Brockman et al., 2010; Custodio et al., 2013), we control for several firm characteristics known to influence firms' decision in using short-term debt including Tobin's Q, leverage, size, size squared, profitability, credit rating, abnormal earnings, tangibility, asset volatility, and asset maturity. Tobin's Q as the proxy for growth opportunities is obtained by dividing the sum of market value of equity plus book value of debt by total assets. Leverage is the ratio of the sum of long-term debt and debt in current liabilities to total assets. Size is the natural log of total assets. We also include size squared to control for the nonlinear relation between credit quality and debt maturity. Profitability is the ratio of earnings before interest, taxes, depreciation, and amortization to total assets. and asset tangibility the ratio of property, plant, and equipment to total assets.

Credit rating is an indicator variable equal to one if the firm has an S&P long-term rating, and zero otherwise. Abnormal earnings are defined as the ratio of difference between the income before extraordinary items, adjusted for common or ordinary stock (capital) equivalents for time t and t-1 to the market value of equity. Tangibility is obtained by dividing property, plant, and equipment by total assets, and asset volatility is defined as the standard deviation of stock return during the fiscal year times the ratio of market value of equity to market value of assets. Asset maturity is calculated as the ratio of property, plant, and equipment to the sum of depreciation and amortization times the proportion of property, plant, and equipment in total assets, plus the ratio of current assets to the cost of goods sold times the proportion of current assets in total assets.

In addition to firm characteristics, we also consider, in line with previous studies (e.g., Li et al., 2018; Qiu and Wang, 2018), state-level economic and political variables that capture the effect of the local environment on firms' debt maturity decisions. We account for the effect of state political environment on firm debt structure by following Klasa et al. (2018) in constructing a proxy for state political balance, defined as the fraction of Congress members representing a given state in the U. S. House of Representatives that belong to the Democratic Party. This data is collected from the U.S. House of Representatives. We account for local economic conditions by including states' annual GDP growth rates, which are collected from the Bureau of Economic Analysis (BEA).

3.4. Descriptive statistics

Table 2 presents summary statistics for our sample. The average

[2] Our sample starts from 1977 following Klasa et al. (2018).

[3] Firms' historical headquarter state data from 2003 is obtained from the SEC EDGAR filings. Bai et al. (2020) provide the data from 1969 to 2003 and Gao et al. (2021) the data from 1994 to 2018. We thank the authors for making the data available at https://sites.google.com/utk.edu/matthew-serfling/research and https://mingze-gao.com/posts/firm-historical-headquarter-state-from-10k/#how-to-get-the-actual-historical-firm-hq-state-using-sec-filings, respectively.

[4] We thank Buhui Qiu for sharing the data on states adopting or rejecting the IDD. The data is available at https://sites.google.com/site/buhuiqiu/research.

Table 1
States adopting or rejecting the IDD.

State	Cases	Date	State court decision
Arkansas	Southwestern Energy v. Eickenhorst, 955 F. Supp. 1078 (1997)	March 18, 1997	Adoption
California	Electro Optical Indus., Inc. v. Stephen White, 90 Cal. Rptr. 2d 680 (1999), 76 Cal. App. 4th 653	November 30, 1999	Adoption
California	Whyte v. Schlage Lock Co., No. G028382 (Ct. of App. of California 2002)	September 12, 2002	Rejection
Connecticut	Branson Ultrasonics Corp. v. Stratman, 921 F. Supp. 909 (D. Conn. 1996)	February 28, 1996	Adoption
Delaware	E.I. DuPont de Nemours & Co. v. American Potash and Chemical Corp., 200 A. 2d 428 (Del Ch. 1964)	May 5, 1964	Adoption
Florida	Fountain v. Hudson Cush-N-Foam Corp., 122 So. 2d 232, 234 (Fla. Dist. Ct. App. 1960)	July 11, 1960	Adoption
Florida	Del Monte Fresh Produce Co. v. Dole Food Co., 148 F. Supp. 2d 1326 (S.D. Fla. 2001)	May 24, 2001	Rejection
Iowa	Barilla Am., Inc. v. Wright, No. 4–02-CV-90267, 2002 U.S. Dist. Lexis 12773 (S.D. Iowa 2002)	July 5, 2002	Adoption
Illinois	PepsiCo, Inc. v. Redmond, 54 F.3d 1262, 1272 (7th Cir. 1995)	May 11, 1995	Adoption
Indiana	Ackerman v. Kimball Int'l, Inc., 652 N.E.2d 507, 510–11 (Ind. 1995)	July 12, 1995	Adoption
Indiana	Bridgestone/Firestone, Inc. v. Lockhart, 5 F. Supp. 2d 667 (S.D. Ind. 1998)	May 7, 1998	Rejection
Kansas	Bradbury Co. v. Teissier-Ducros, 413 F. Supp. 2d 1203, 1209 (D. Kan. 2006)	February 2, 2006	Adoption
Massachusetts	Marcam Corp. v. Orchard, 885 F. Supp. 294, 298–300 (D. Mass. 1995)	April 3, 1995	Adoption
Massachusetts	U.S. Elec. Servs. v. Schmidt, Civil Action No. 12–10845-DJC (U.S. Dist. CT. for the Dist. of Mass. 2012)	June 19, 2012	Rejection
Michigan	Allis-Chalmers Manufacturing Co. v. Continental Aviation & Engineering Corp., 255 F. Supp. 645, 654 (E.D. Mich. 1966)	February 17, 1966	Adoption
Michigan	CMI International Inc. v. Intermet Inter. Corp., 649 N.W.2d 808 (Mich. Ct. App.2002)	April 30, 2002	Rejection
Minnesota	Surgidev Corp. v. Eye Tech., Inc., 648 F. Supp. 661 (D. Minn. 1986)	October 10, 1986	Adoption
Minnesota	IBM Corp. v. Seagate Tech., Inc., 941 F. Supp. 98 (D. Minn. 1992)	April 21, 1992	Rejection
Minnesota	La Calhene, Inc. v. Spolyar, 938 F. Supp. 523 (W.D. Wis. 1996)	August 23, 1996	Adoption
Missouri	H&R Block Eastern Tax Services, Inc. v. Enchura, 122 F. Supp. 2d 1067 (W.D.Mo. 2000)	November 2, 2000	Adoption
New Jersey	National Starch and Chem. Corp. v. Parker Chemical Corp., 530 A.2d 31 (N.J. Super. Ct. App. Div. 1987)	April 27, 1987	Adoption
New York	DoubleClick, Inc. v. Henderson, No. 116914/97, 1997 N.Y. Misc. Lexis 577 (Sup. Ct. N.Y. Co. Nov. 7, 1997)	November 7, 1997	Adoption
New York	EarthWeb, Inc. v. Schlack, 71 F. Supp. 2d 299 (S.D.N.Y. 1999)	October 27, 1999	Rejection
North Carolina	Travenol Labs., Inc. v. Turner, 228 S.E.2d 478, 483 (N.C. Ct. App. 1976)	October 6, 1976	Adoption
North Carolina	RCR Enters., LLC v. McCall, 14 CVS 3342 (N.C. Sup. Ct. 2014)	December 19, 2014	Rejection
Ohio	Procter & Gamble Co., v. Stoneham, 747 N.E.2d 268 (Ohio Ct. App. 2000)	September 29, 2000	Adoption
Pennsylvania	Air Products & Chemical, Inc. v. Johnson, 442 A.2d 1114 (Pennsylvania Superior Ct. 1982)	February 19, 1982	Adoption
South Carolina	Nucor Corp. v. Bell, C/A No. 2: 06-CV-02972-DCN (U.S. Dist. Ct. for the Dist. of South Carolina 2008)	March 14, 2008	Adoption
Texas	Rugen v. Interactive Bus. Sys., Inc., 864 S.W.2d 548, 551 (Tex. App. 1993)	May 28, 1993	Adoption
Texas	Cardinal Health Staffing Network Inc. v. Bowen, 106 S.W.3d 230 (Tex. App. 2003)	April 3, 2003	Rejection
Utah	Novell, Inc. v. Timpanogos Research Group, Inc., 46 U.S.P.Q.2d 1197 (Utah Dist. Ct. 1998)	January 30,1998	Adoption
Washington	Solutec Corp., Inc. v. Agnew, 1997 WL 794496, 8 (Wash. Ct. App.)	December 30, 1997	Adoption

Note: This table lists the legal cases in which state courts adopted the IDD or rejected it after adopting it from 1960 to 2015, which are reproduced from Table 1 of Qiu and Wang (2018).

Table 2
Descriptive statistics.

Variables	Mean	Median	Std. Dev.	P25	P75	Obs.
LT1	0.720	0.855	0.315	0.571	0.965	90,140
LT2	0.583	0.691	0.350	0.271	0.888	76,763
LT3	0.470	0.516	0.354	0.069	0.788	76,620
LT4	0.376	0.350	0.339	0.001	0.669	76,401
LT5	0.289	0.186	0.310	0	0.536	75,126
IDD	0.321	0	0.467	0	1	75,126
Tobin's Q	1.821	1.351	1.802	1.038	1.956	75,126
Leverage	0.258	0.228	0.216	0.101	0.364	75,126
Size	5.409	5.249	2.021	3.902	6.781	75,126
Size squared	33.340	27.554	23.736	15.224	45.986	75,126
Profitability	0.084	0.122	0.232	0.063	0.176	75,126
Credit rating	0.249	0	0.433	0	0	75,126
Abnormal earnings	-0.013	0.007	0.354	-0.032	0.032	75,126
Tangibility	0.562	0.488	0.376	0.278	0.778	75,126
Asset volatility	0.023	0.018	0.020	0.011	0.029	75,126
Asset maturity	9.443	6.691	9.893	3.406	12.242	75,126
Political balance (%)	55.856	56.667	18.867	46.154	64.151	75,126
State GDP growth (%)	5.893	5.703	3.257	3.881	7.795	75,126

Note: This table reports the descriptive statistics for variables used in the baseline regressions from 1977 to 2015. Variable definitions are provided in Appendix A. All the continuous variables are winsorized at the 1% and 99% levels.

(median) of *LT5* equals 0.289 (0.186). The average (median) of *LT3* is 0.47 (0.516). The numbers for average and median of debt maturity measures are comparable to those reported in Custodio et al. (2013). The mean value of *IDD* is 0.321, indicating that 32.1% of firm-year observations are in states with the IDD adoption. An average firm in our sample has firm size of 5.409, Tobin's Q of 1.821, leverage ratio of 0.258, profitability of 0.084, abnormal earnings of − 0.013, tangibility of 0.562, asset volatility of 0.023, and asset maturity of 9.443. The mean value of credit rating is 0.249, suggesting that 24.9% of firm-year observations have a credit rating.

4. Empirical model and results

4.1. Baseline estimation

To examine the causal relation between labor mobility restrictions and debt maturity, we employ state-level shocks as natural experiments. We implement the DiD approach by assigning firms headquartered in states that have (have not) adopted the IDD as the treatment (control) group. The staggered adoption of the IDD enables us to compare firms headquartered in different states and subject to different regulations. Following Bertrand and Mullainathan (2003), we specify our DiD model as follows:

$$Debtmaturity_{it} = \beta_0 + \beta_1 IDD_{it-1} + \boldsymbol{\beta_2} X_{it-1} + \gamma_i + \delta_j + \varepsilon_{it} \quad (1)$$

where i indexes firm, j indexes industry-year, and t indexes year. *Debt maturity*$_{it}$ is the *LT5* of firm i in year t. IDD_{it-1} is an indicator variable that equals one if firm i is headquartered in a state that has adopted the IDD from year *t-1*, and zero otherwise. The coefficient β_1 captures the change in debt maturity for firms headquartered in IDD-adopting states relative to the debt maturity for firms headquartered in non-IDD-adopting states. $X_{it-1}$is a set of control variables consisting of lagged value of Tobin's Q, leverage, firm size, firm size squared, profitability, credit rating, abnormal earnings, tangibility, asset volatility, asset maturity, and state-level controls including state GDP growth and political balance. $\gamma_i$is firm fixed effects that capture time-invariant unobservable firm characteristics. δ_j is industry-year fixed effects that enable us to control for time-varying industry heterogeneity. $\boldsymbol{\varepsilon}_{it}$is the error term. Since the IDD_{it-1} indicator is at the state level, standard errors are clustered by the states in which firms are headquartered. The variable definitions are provided in Appendix A.

Table 3 reports the DiD regression results with debt maturity (*LT5*) as the dependent variable. We exclude firm and year fixed effects in column 1. A few fixed effects are included in columns 2–4. We include firm fixed effects in column 2, firm and year fixed effects in column 3, and firm and industry-year fixed effects in column 4 (Eq. (1)). The coefficients of IDD_{it-1} are negative and statistically significant in all four columns.[5] These results lend support to our hypothesis *H1b* that firms headquartered in IDD-adopting states tend to use short-term debt financing following adoption of the IDD, relative to non-IDD adopting states. In terms of the economic magnitude, our estimations reveal that the long-term debt usage decreases by 1.2% after IDD adoption, which is equivalent to 4.15% (6.45%) of the average (median) *LT5* in our sample.[6]

4.2. Alternative measures of debt maturity

To ensure that the baseline regression results are not sensitive to the way the debt maturity is defined, apart from using *LT5* as the measure of debt maturity, we repeat the baseline regression using four alternative measures of debt maturity, *LT1* to *LT4*. *LT1* is the proportion of long-term debt in total debt, while *LT2*, *LT3*, and *LT4* are, respectively, the proportion of long-term debt maturing more than two, three, and four years. Table 4 reports the regression results using these four measures as well as the one for *LT5* for comparison. As expected, the coefficient of IDD_{it-1} is negative and statistically insignificant when the dependent variable is *LT1* and *LT2*, respectively (i.e., columns 1 and 2 in Table 4), since the numerator of these two measures include debts that will mature within one or two years. The coefficient of IDD_{it-1} is negative and statistically significant when the dependent variable is *LT3* and *LT4*, respectively (i.e., columns 3 and 4 in Table 4). That our results are consistent with the results based on *LT5* suggests that firms are more inclined to use short-term debt with a lower proportion of debt maturing more than 3 years in the wake of the IDD adoption.

4.3. Validity of the identification strategy

The main potential endogeneity concern for our baseline results is that our results may be driven by simultaneity or unobservable characteristics. This happens if the change in a firm's debt maturity structure precedes adoption/rejection of the IDD. In other words, the assumption behind the DiD research design that the trends in the debt maturity of the treatment firms in adopting states and control firms in non-adopting states are parallel may not hold. The parallel trends assumption is critical to ensuring the validity of the DiD estimates, but inherently untestable. To address this concern, we perform a few robustness checks to ensure that our results are not driven by non-parallel trends, other unobservable characteristics, or simultaneity.

First, we conduct a placebo test by creating placebo *IDD* indicator variables equal to one for fictitious changes in the IDD that occur one and two years before and after the year when a state adopted/rejected the IDD and repeat the baseline analysis (Qiu and Wang, 2018). The rationale for the test is if our results are driven by unobservable characteristics that simultaneously affect debt maturity and IDD adoption/rejection, then the coefficient of the placebo IDD indicator variables for fictitious changes in the IDD that occur one and two years before the actual IDD event should be statistically significant. This would imply that the parallel trends assumption may fail to hold, thus casting doubt on the validity of our empirical approach. Panel A of Table 5 reports the results of re-running the DiD regression for our baseline model (Eq. 1) using these placebo *IDD* indicator variables. Columns 1–5 show the coefficients of the placebo *IDD* indicators to be negative in all regressions, but statistically insignificant for the placebo indicators of one and two years before the actual IDD and two years after the actual IDD.[7]

Second, we explore the dynamic effect of the IDD adoption on debt maturity. To this end, we follow Klasa et al. (2018) and Li et al. (2018) by investigating the timing of changes in debt maturity relative to the timing of adoptions of the IDD. The key variables of interest are *IDD (−2)*, *IDD (−1)*, *IDD (0)*, *IDD (+1)*, and *IDD (2 +)*, which are equal to one if the state court will adopt the IDD in two years, will adopt the IDD in one year, adopts the IDD in the current year, adopted the IDD one year ago, and adopted the IDD two or more years ago, respectively, and zero

Table 3
Baseline regression.

Variables	(1) LT5	(2) LT5	(3) LT5	(4) LT5
IDD_{it-1}	-0.029 * **	-0.029 * **	-0.009 *	-0.012 * *
	(−2.735)	(−4.259)	(−1.920)	(−2.019)
Tobin's Q_{it-1}	-0.003 * *	0.000	0.002 * *	0.002 * *
	(−2.386)	(0.435)	(2.638)	(2.130)
$Leverage_{it-1}$	0.147 * **	0.091 * **	0.083 * **	0.081 * **
	(9.608)	(7.888)	(6.917)	(6.462)
$Size_{it-1}$	0.046 * **	-0.007	0.009	0.012
	(10.262)	(−0.938)	(1.170)	(1.278)
Size $squared_{it-1}$	-0.001 * **	0.001	0.002 * **	0.001 *
	(−3.424)	(1.469)	(2.997)	(1.781)
$Profitability_{it-1}$	0.051 * **	0.038 * **	0.019 * **	0.013 *
	(8.308)	(5.760)	(2.797)	(1.896)
Credit $rating_{it-1}$	0.152 * **	0.079 * **	0.090 * **	0.088 * **
	(16.152)	(9.652)	(11.610)	(11.420)
Abnormal $earnings_{it-1}$	0.015 * **	0.007 * **	0.009 * **	0.009 * **
	(6.729)	(4.023)	(4.870)	(4.586)
$Tangibility_{it-1}$	0.019 * *	-0.080 * **	-0.048 * **	-0.048 * **
	(2.532)	(−7.710)	(−5.002)	(−4.982)
Asset $volatility_{it-1}$	-1.076 * **	-0.184 * **	0.061	0.099
	(−10.382)	(−2.826)	(0.625)	(0.975)
Asset $maturity_{it-1}$	0.002 * **	0.001 * **	0.001 * **	0.001 * **
	(5.132)	(4.544)	(5.266)	(4.027)
State GDP $growth_{it-1}$	0.015 * **	0.007 * **	0.002 * **	0.001 * *
	(12.088)	(11.390)	(3.469)	(2.335)
Political $balance_{it-1}$	0.000 *	0.000	0.000	0.000
	(1.830)	(1.055)	(0.512)	(0.758)
Constant	-0.101 * **	0.250 * **	0.131 * **	0.133 * **
	(−5.730)	(9.918)	(5.098)	(4.902)
Firm FE	N	Y	Y	Y
Year FE	N	N	Y	N
Ind*Year FE	N	N	N	Y
Observations	76,455	75,214	75,214	75,126
Adjusted R-squared	0.240	0.482	0.491	0.497

Note: This table reports the results from regressing debt maturity (*LT5*) on the *IDD* indicator variable and control variables. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * ** , * *, and * represent significance at the 1st, 5th, and 10th percentile levels, respectively.

[5] Our results remain qualitatively unchanged if we exclude California from the analysis or set the IDD indicator to be 0 for California or set the IDD indicator to be 0.5 for California from 1999 to 2001 following Qiu and Wang (2018).

[6] 4.15% (6.45%) is calculated as 1.2%/0.289 (1.2%/0.186).

[7] Because we use a lagged one-year *IDD* indicator in our regressions, the statistically significant result for actual IDD change year indicates that the IDD changes may have some impact in the year in which the IDD is adopted or rejected.

Table 4
Alternative measures of debt maturity.

Variables	(1) LT1	(2) LT2	(3) LT3	(4) LT4	(5) LT5
IDD_{it-1}	-0.008	-0.009	-0.010 *	-0.011 * *	-0.012 * *
	(−1.300)	(−1.390)	(−1.780)	(−2.034)	(−2.019)
Tobin's Q_{it-1}	0.003 * **	0.003 * **	0.004 * **	0.004 * **	0.002 * *
	(2.915)	(3.151)	(3.068)	(3.788)	(2.130)
$Leverage_{it-1}$	0.139 * **	0.171 * **	0.147 * **	0.077 * **	0.081 * **
	(8.463)	(8.471)	(8.687)	(5.858)	(6.462)
$Size_{it-1}$	0.032 * **	0.055 * **	0.055 * **	0.038 * **	0.012
	(3.356)	(4.315)	(4.666)	(3.889)	(1.278)
Size $squared_{it-1}$	-0.000	-0.001	-0.001	0.000	0.001 *
	(−0.102)	(−0.668)	(−0.549)	(0.164)	(1.781)
$Profitability_{it-1}$	0.047 * **	0.029 * *	0.023 * *	0.015 *	0.013 *
	(3.717)	(2.107)	(2.167)	(1.740)	(1.896)
Credit $rating_{it-1}$	0.035 * **	0.059 * **	0.083 * **	0.093 * **	0.088 * **
	(5.479)	(10.169)	(12.654)	(13.547)	(11.420)
Abnormal $earnings_{it-1}$	0.016 * **	0.015 * **	0.010 * **	0.011 * **	0.009 * **
	(6.013)	(4.226)	(3.191)	(4.173)	(4.586)
$Tangibility_{it-1}$	-0.038 * **	-0.050 * **	-0.047 * **	-0.039 * **	-0.048 * **
	(−3.772)	(−5.068)	(−5.552)	(−4.596)	(−4.982)
Asset $volatility_{it-1}$	0.158 *	0.184 *	0.272 * **	0.184 *	0.099
	(1.908)	(1.854)	(2.817)	(1.732)	(0.975)
Asset $maturity_{it-1}$	0.000	0.000 *	0.001 * **	0.001 * **	0.001 * **
	(0.489)	(1.768)	(3.098)	(3.042)	(4.027)
State GDP $growth_{it-1}$	0.001	0.000	0.001 *	0.001 *	0.001 * *
	(1.442)	(0.837)	(1.908)	(1.829)	(2.335)
Political $balance_{it-1}$	-0.000	0.000	0.000	0.000	0.000
	(−0.624)	(0.182)	(0.992)	(0.054)	(0.758)
Constant	0.520 * **	0.263 * **	0.125 * **	0.123 * **	0.133 * **
	(17.541)	(7.876)	(3.760)	(4.414)	(4.902)
Firm FE	Y	Y	Y	Y	Y
Ind*Year FE	Y	Y	Y	Y	Y
Observations	90,140	76,763	76,620	76,401	75,126
Adjusted R-squared	0.453	0.498	0.504	0.485	0.497

Note: This table reports the coefficient estimates of the *IDD* indicator variable using alternative measures of firm's debt maturity as dependent variable. Columns 1–5 use *LT1* to *LT5* as the dependent variable, respectively. We include the results of *LT5* for comparison purpose. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * ** , * *, and * represent significance at the 1st, 5th, and 10th percentile levels, respectively.

otherwise. Again, statistically significant coefficients on *IDD (−2)* and *IDD (−1)* imply the existence of pre-treatment trends, hence the violation of parallel trends assumption. The regression results are reported in Panel B of Table 5. We find that the coefficients on *IDD (−2)* and *IDD (−1)* are close to zero and statistically insignificant, while the coefficients on *IDD (0)* and *IDD (+1)* are negative and statistically significant. These suggest that our results are unlikely to be driven by trend differences between the treatment and control groups. In other words, the decrease in the long-term debt financing for treatment firms in IDD adopting states relative to control firms in non-adopting states occurs after adoption of the IDD, but not before.

Third, we plot the difference in average debt maturity surrounding the IDD adoption between firms headquartered in states with the IDD adoption and those without. Fig. 1 shows that the difference is relatively stable before the IDD adoption and starts to drop after the IDD adoption, which is consistent with *H1b* that the IDD adoption is positively associated with short-term debt financing. This provides graphical evidence to show that parallel trends assumption is not violated.

Fourth, we estimate the effect of IDD adoption on debt maturity using propensity score matched samples to address selection bias. The existence of the parallel trends pre-treatment is based on the assumption that the debt maturity structure of the treatment and control firms are not systematically different without the IDD adoption. For this matching procedure, each firm in an IDD-adopting state (i.e., treatment group) is matched with a control firm in a non-IDD-adopting state (i.e., control group)[8] based on the propensity score. We run a probit regression with the dependent variable equals one if the firm belongs to the treatment group, and zero if it belongs to the control group. The probit regression includes all firm-level control variables in Eq. (1).

Column 1 of Panel A of Table 6 reports the results of the probit model for the propensity-score-matching analysis, while column 2 the results after performing the nearest-neighbor propensity-score matching without replacement. Comparing those two columns, we find that 1) none of the independent variables is statistically significant post-matching; 2) the pseudo-R^2 drops from 1.37% pre-matching to 0.3% post-matching; and 3) the p-value of the Chi-squared test is 0.494 post-matching. These results suggest that the null hypothesis of all coefficients are zero cannot be rejected. Overall, the results in Panel A indicate that the parallel-trends assumption is not violated. Next, we re-estimate the baseline regressions using the observations three years before and after the IDD adoption. The coefficient of IDD_{it-1} reported in Panel B is negative and statistically significant with or without fixed effects, which is consistent with our previous findings that firms headquartered in IDD-adopting states tend to use short-term debt following adoption of the IDD.

Fifth, although the staggered DiD methodology has been widely used in accounting and finance studies for causal interpretation, recent literature raised a concern that different timing of treatment of different units (e.g., states) leads to biases in traditional estimation of the Average Treatment of Treated group (ATT) (Barrios, 2021; Callaway and Sant'Anna, 2021; Goodman-Bacon, 2021; Baker et al., 2022). To alleviate this concern, we re-run the main regression using a stacked regression approach following Baker et al. (2022). Specifically, by stacking and aligning events in event-time, this approach prevents using past treated units as effective comparison units, which is equivalent to a setting with events happening contemporaneously (Cengiz et al., 2019; Aswani et al., 2021; Baker et al., 2022). The results are reported in Table 7. The

[8] Control firms are firms from states that do not experience IDD adoption throughout a 7-year window (i.e., 3 years before and after) following Bourveau et al. (2018).

Table 5
The IDD and the timing of debt maturity changes.

Panel A: Placebo tests					
	(1)	(2)	(3)	(4)	(5)
Variables	Placebo change 2 years before the actual change	Placebo change 1 year before the actual change	Actual IDD change	Placebo change 1 year after the actual change	Placebo change 2 years after the actual change
IDD /Placebo IDD	-0.010	-0.011	-0.012 * *	-0.012 * *	-0.007
	(−1.565)	(−1.539)	(−2.039)	(−2.019)	(−1.208)
Controls	Y	Y	Y	Y	Y
Firm FE	Y	Y	Y	Y	Y
Ind*Year FE	Y	Y	Y	Y	Y
Observations	71,775	73,433	75,126	75,126	75,126
Adjusted R-squared	0.502	0.500	0.497	0.497	0.497

Panel B: Dynamic effects	
Variables	LT5
IDD(−2)	-0.006
	(−0.934)
IDD(−1)	-0.007
	(−0.902)
IDD(0)	-0.015 * *
	(−2.154)
IDD(+1)	-0.022 * **
	(−2.737)
IDD(2 +)	-0.013
	(−1.610)
IDD $\text{Rejection}_{it\text{-}1}$	0.008 * *
	(2.125)
Controls	Y
Firm FE	Y
Ind*Year FE	Y
Observations	75,126
Adjusted R-squared	0.497

Note: This table reports the results of placebo tests regarding the effects of actual and fictitious changes in the IDD on debt maturity and the dynamic effect of the IDD on debt maturity. We control for lagged firm characteristics, state variables, firm fixed effects, and industry-by-year fixed effects in all regressions. For the placebo experiments in Panel A, we follow Qiu and Wang (2018) to create fictitious changes in the IDD that take place one and two years before and after the actual adoption/rejection year of the IDD in a state. In Panel B, following Klasa et al. (2018) and Li et al. (2018), we define IDD(−2), IDD(−1), and IDD(0) as indicator variables that equal one if the state court will adopt the IDD in 2 years, 1 year, and by the end of the current year, respectively. IDD(+1) is an indicator variable that equals one if the state court adopted the IDD one year ago, and zero otherwise. IDD(2 +) is an indicator variable that equals one if the state court adopted the IDD two or more years ago, and zero otherwise. *IDD Rejection* is an indicator variable that equals one if the state where the firm is headquartered rejects the previously adopted IDD, and zero otherwise. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * ** and * * represent significance at the 1st and 5th percentile levels, respectively.

coefficient of IDD_{it-1} is negative and statistically significant at the 5% level, which confirms that the statistical significance of our results is unaffected by this bias.

Furthermore, we conduct two additional tests to allay the concern that the IDD changes may not be fully exogenous, but occur because of local shocks. Sixth, to account for the potential effect of state legal conditions on firms' debt maturity and mitigate estimation bias caused by omitted variables, we include additional state legal control variables, namely, an indicator variable for Right-to-Work, and an indicator variable that equals one if a state has adopted the UTSA. The first variable,

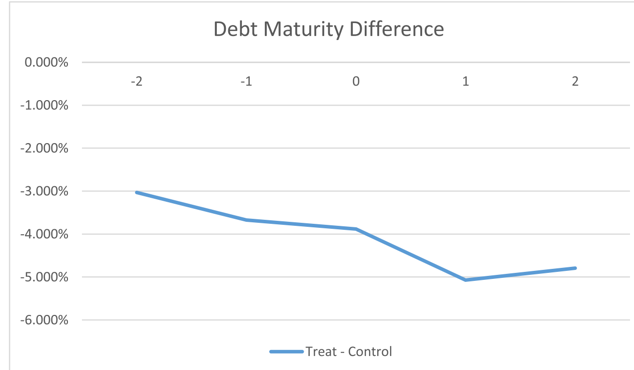

Fig. 1. Difference in average debt maturity in response to the IDD adoption, This figure shows the difference in average debt maturity between treatment and control firms around the IDD adoption years. Treatment firms are firms headquartered in states where state courts adopted the IDD and control firms are those headquartered in states without the IDD change.

RtW, equals one for states that have passed Right-to-Work legislation, and zero otherwise. In states in which Right-to-Work legislation has been passed, labor unions are prohibited from enforcing compulsory union membership and payment of the union dues, which prohibitions substantially reduce union bargaining power (Ellwood and Fine, 1987). This variable is included to mitigate the concern that union bargaining power may drive adoption of the IDD and the increase in short-term debt financing. The second, *UTSA*, equals one if a firm is headquartered in a state in which UTSA is enacted, and zero otherwise. Being one of the legal tools that afford trade secret protection, we include UTSA to mitigate the concern that its enactment is driving our results. Column 1 (2) of Table 8 presents the DiD regression results, which include *RtW* (*UTSA*), and in column 3 we augment the baseline model by including both indicator variables. The results confirm hypothesis *H1b*, which maintains that firms headquartered in IDD-adopting states rely more heavily on short-term debt following the IDD adoption.

Seventh, to further alleviate the concern that unobserved confounding factors associated with adoption of the IDD and firms' debt maturity may result in spurious relations, we investigate in an additional test the reaction of firms headquartered in states neighboring the states that experienced the IDD change. The rationale for the tests is the idea that time-varying local market dynamics may spill across state borders and affect firms located in contiguous states. If IDD changes and firms' debt maturity decisions are driven by unobservable time-varying local dynamics, the IDD changes in a given state will lead firms in neighboring states to hold more short-term debt. Following Qiu and Wang (2018), we conduct a neighboring-state test by creating a placebo *IDD* indicator variable equal to one for the occurrence of fictitious changes in the IDD that occur in the year when neighboring states adopted/rejected the IDD. Next, similar to Panel B of Table 5, we also create the variables $IDD_{neighboring}\,(-2)$, $IDD_{neighboring}\,(-1)$, $IDD_{neighboring}\,(0)$, $IDD_{neighboring}\,(+1)$, and $IDD_{neighboring}\,(2\,+)$, which are equal to one if the neighboring-state court will adopt the IDD in two years, will adopt the IDD in one year, adopts the IDD in the current year, adopted the IDD one year ago, and adopted the IDD two or more years ago, respectively, and zero otherwise (Li et al., 2018). We then re-run the baseline regression and the dynamic analysis using the placebo *IDD* indicator variables. As shown in columns 1 and 2 of Table 9, none of the coefficients of the placebo *IDD* indicator variables is statistically significant, which confirms our baseline findings and alleviates the concern that our findings are driven by unobservable time-varying local dynamics.

Taken together, the findings from the foregoing robustness checks to allay the endogeneity concerns continue to support our results that labor mobility restrictions are positively associated with short-term debt financing.

Table 6
Propensity Score Matching Analysis.

Panel A: Pre-match propensity-score regression and post-match diagnostic regression

	Indicator = 1 if in treatment group, 0 if in control group	
Variables	(1) Pre-match	(2) Post-match
Tobin's Q_{it-1}	0.013 * **	0.003
	(2.815)	(0.172)
$Leverage_{it-1}$	-0.074	-0.261
	(−0.867)	(−1.493)
$Size_{it-1}$	-0.044	-0.004
	(−0.739)	(−0.030)
Size $squared_{it-1}$	-0.001	0.002
	(−0.347)	(0.215)
$Profitability_{it-1}$	0.089	-0.113
	(1.104)	(−0.707)
Credit $rating_{it-1}$	0.214 * **	0.048
	(3.710)	(0.499)
Abnormal $earnings_{it-1}$	-0.007	0.077
	(−0.255)	(0.899)
$Tangibility_{it-1}$	0.020	-0.081
	(0.291)	(−0.520)
Asset $volatility_{it-1}$	3.867 * *	-0.863
	(2.343)	(−0.145)
Asset $maturity_{it-1}$	-0.005 *	-0.002
	(−1.953)	(−0.662)
Constant	-1.545 * **	0.102
	(−6.292)	(0.164)
Observations	58,149	5,498
Pseudo R-squared	0.0137	0.003
P-value of Chi-squared	0	0.494

Panel B: Regression Analysis

Variables	(1) LT5	(2) LT5
IDD $_{it-1}$	-0.025 * *	-0.010 *
	(−2.552)	(−1.694)
Controls	Y	Y
Firm FE	N	Y
Ind*Year FE	N	Y
Observations	17,994	17,346
Adjusted R-squared	0.245	0.595

Note: This table reports the results from regressing debt maturity on the IDD changes, using propensity score matched samples. Panel A reports coefficient estimates from the probit model employed in estimating the propensity scores for firms in the treatment and control groups. The dependent variable is an indicator variable that equals one if the firm is headquartered in the state that adopts the IDD in the year, and zero otherwise. Panel B reports the coefficient estimates for the change in debt maturity of treatment and control firms surrounding the IDD change. Control variables are not reported in Panel B for brevity. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * ** , * *, and * represent significance at the 1st, 5th, and 10th percentile levels, respectively.

Table 7
Stacked regression.

Variables	(1) LT5
IDD_{it-1}	-0.009 * *
	(−2.040)
Controls	Yes
Firm FE	Yes
Year FE	Yes
Observations	76,455
Number of groups	9,827

Note: This table reports the results from regressing debt maturity (*LT5*) on the *IDD* indicator variable and control variables using the stacked regression approach. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * * represents significance at the 5th percentile level.

Table 8
State legal variables.

Variables	(1) LT5	(2) LT5	(3) LT5
IDD_{it-1}	-0.012 * *	-0.012 * *	-0.012 * *
	(−2.052)	(−2.100)	(−2.127)
RtW_{it-1}	-0.013		-0.013
	(−1.369)		(−1.338)
$UTSA_{it-1}$		-0.005	-0.005
		(−0.781)	(−0.806)
Controls	Y	Y	Y
Firm FE	Y	Y	Y
Ind*Year FE	Y	Y	Y
Observations	75,126	75,126	75,126
Adjusted R-squared	0.497	0.497	0.497

Note: This table reports the results from regressing debt maturity on the IDD changes, controlling for other state-level laws: RtW and UTSA. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * * represents significance at the 5th percentile level.

5. Mechanisms analyses

Our results thus far provide robust evidence consistent with hypothesis *H1b* that firms tend to increase short-term debt usage after an increase in labor mobility restrictions. We now explore the three non-mutually exclusive mechanisms through which labor mobility restrictions affect debt maturity via the IDD adoption: 1) default risk, 2) information asymmetry, and 3) agency costs mitigation.

5.1. Default risk mechanism

Labor mobility restrictions can lower firms' default risk in several ways. First, they reduce the competitive threats by limiting the leakage of trade secrets to the rival firms. Second, since labor mobility restrictions make poaching employees difficult, they reduce the cost both of hiring labor and debt (Png and Samila, 2013; Choi, 2020). Third, by reducing managers' outside employment options, labor mobility

Table 9
Placebo tests.

Variables	(1) LT5	(2) LT5
$IDD_{neighboring}$	0.003	
	(0.438)	
$IDD_{neighboring}$ (−2)		0.005
		(1.356)
$IDD_{neighboring}$ (−1)		-0.005
		(−0.752)
$IDD_{neighboring}$ (0)		-0.001
		(−0.226)
$IDD_{neighboring}$ (+1)		-0.006
		(−1.345)
$IDD_{neighboring}$ (2 +)		0.004
		(0.577)
$IDD_{neighboring}$ $Rejection_{it-1}$		0.004
		(0.845)
Controls	Y	Y
Firm FE	Y	Y
Ind*Year FE	Y	Y
Observations	75,126	75,126
Adjusted R-squared	0.497	0.497

Note: This table reports the results of placebo tests regarding the effects of fictitious changes in the IDD on debt maturity. In column 1 we follow Qiu and Wang (2018) in examining the effect of neighboring state IDD adoption and rejection on firms' debt maturity. Similar to Panel B of Table 5, in column 2 we create fictitious changes in the IDD that take place in the actual adoption/rejection year of the IDD in the neighboring states for the placebo experiments. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * ** , * *, * represent significance at the 1st, 5th, and 10th percentile levels, respectively.

restrictions help prevent managerial myopia by motivating managers to focus on long-term performance rather than short-term performance (Islam et al., 2021) and reducing managerial risk-taking incentives (Islam et al., 2019; Cici et al., forthcoming). Since a lower default risk reduces the concern of liquidity risk, firms may prefer short-term debt to long-term debt.

To test this prediction, we employ two default risk proxies: Altman's (1968) z-score and KZ index (Kaplan and Zingales, 1997). Both measures are accounting-based measures of the financial health of a company. A higher (lower) z-score (KZ index) indicates low default risk. We include these two proxies and their respective interactions with IDD_{it-1} in our baseline model. As shown in columns 1 and 2 of Table 10, the interaction term coefficients are not statistically significant. The evidence shows that default risk does not explain the positive relation between labor mobility restrictions and short-term debt financing.

5.2. Information asymmetry mechanism

Previous studies suggest that increased labor mobility restrictions constrain the flow of information from firms to rivals, resulting in an increase in the cost of information disclosure (e.g., Aobdia, 2018; Li et al., 2018; Callen et al., 2020). Firms consequently become more opaque and experience more severe information asymmetry. Firms in such deteriorating information environments may exhibit a preference for short-term debt because it is less sensitive to shifts in risk (Barnea et al., 1980). Consequently, we expect restricting labor mobility via the IDD adoption to result in an increase in information asymmetry, leading firms to prefer the more intensive monitoring provided by short-term debt.

To test this prediction, we augment the baseline regression in Eq. (1) by including proxies for information asymmetry and their interactions with the *IDD* indicator variable. We use analyst forecast dispersion and analyst forecast error as proxies for degree of information asymmetry (Glaeser, 2018). A larger analyst forecast dispersion and analyst forecast error indicate higher degree of information asymmetry. Analyst forecast dispersion is calculated as the standard deviation of earnings forecast divided by the previous fiscal-year-end stock price. Analyst forecast error is defined as the absolute value of median forecast errors scaled by the previous fiscal-year-end stock price. Based on the arguments outlined above, we expect the interaction term between the proxies for information asymmetry and IDD_{it-1} to load negatively in our estimations. The results in columns 1 and 2 of Table 11 show the coefficients on the interaction terms to be negative and statistically significant at the 5% level for analyst forecast dispersion, and at the 10% level for analyst forecast error. These results suggest that information asymmetry amplifies the positive relation between labor mobility restrictions and short-term debt financing, which supports our argument that information asymmetry may be the mechanism through which labor mobility restrictions result in the shortening of debt maturity.

Diamond's (1991) findings suggest that the relation between default risk and debt maturity is not monotone. Firms with high or low level of default risk tend to use short-term debt, while those with moderate level of default risk tend to use long-term debt. In addition, reliance too much on short-term debt will expose firms to greater liquidity risk (Liu et al., 2021). To investigate the non-monotonic relationship between default risk and debt maturity, we complement and extend the analysis by splitting the full sample based on two default risk measures, the Altman's (1968) z-score and the KZ index. Specifically, we create two subsamples: 1) the first and fourth quartiles of the z-score (KZ index) equal to one for high (low) and low (high) default risk firms and 2) the second and third quartiles equal to zero for firms with moderate level of default risk. Then we re-run the regressions with the interaction terms between the proxies for information asymmetry and the IDD indicator variable using these two subsamples.

Columns 3–10 of Table 11 report the regression results. Columns 3, 5, 7, and 9 report the results for the low and high default risk subsample and columns 4, 6, 8, and 10 the results for the moderate default risk subsample. The interaction term coefficients are negative and statistically significant for both proxies for information asymmetry in columns 3, 5, 7, and 9. In other words, the effect of information asymmetry on debt maturity is amplified in low and high default risk subsample only. Our results thus provide some evidence to support Diamond's (1991) finding of differential effect of default risk on debt maturity.

5.3. Agency costs mitigation mechanism

An increasing trade secret protection can, by restricting labor mobility, reduce the intensity of product market competition (Klasa et al., 2018). Since product market competition is an effective disciplinary mechanism to mitigate agency problems (e.g., Hart, 1983; Shleifer and Vishny, 1997), firms that face less external monitoring from the product market might thus be expected to exhibit a greater need for the strict monitoring that attends short-term financing. The need for an alternative external governance mechanism could therefore be an underlying economic mechanism through which labor mobility restrictions increase firms' use of short-term debt.

To test this conjecture, we employ as proxies for product market competition, product-market fluidity and TNIC HHI concentration metrics proposed in Hoberg and Phillips (2010, 2016) and Hoberg et al. (2014).[9] Product-market fluidity measures the competitive threats a firm faces in its product market and TNIC HHI concentration metrics a firm's pricing power (Hoberg and Phillips, 2010, 2016). TNIC HHI is positively associated with pricing power. We include in the baseline regression the two proxies for product market competition and their respective interactions with IDD_{it-1}. As can be seen in columns 1 and 2 of Table 12, the coefficients on the interaction terms are positive and statistically insignificant for both. Our results do not support the argument that firms increase short-term debt financing as a governance mechanism to substitute for lower external governance from product market competitive threats.

Table 10
Default risk mechanism.

Variables	(1) LT5	(2) LT5
z-score$_{it-1}$ *IDD$_{it-1}$	0.000	
	(0.829)	
z-score$_{it-1}$	-0.001	
	(−1.415)	
KZ-Index$_{it-1}$ *IDD$_{it-1}$		0.000
		(0.614)
KZ-Index$_{it-1}$		-0.000 *
		(−1.846)
IDD$_{it-1}$	-0.013 * *	-0.011 * *
	(−2.434)	(−2.031)
Controls	Y	Y
Firm FE	Y	Y
Ind*Year FE	Y	Y
Observations	74,442	72,020
Adjusted R-squared	0.496	0.503

Note: This table reports the results of debt maturity regressions augmented with the proxies for default risk. Columns 1 and 2 report the coefficient estimates of the interaction term between the *IDD* indicator variable and two proxies for default risk (i.e., Altman's, 1968 z-score and KZ index). Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * ** , * *, and * represent significance at the 1st, 5th, and 10th percentile levels, respectively.

[9] Product market competition data can be downloaded from http://hoberg-phillips.tuck.dartmouth.edu/tnic_poweruser.htm.

Table 11
Information asymmetry mechanism.

Variables	(1) Full sample	(2) Full sample	(3) Low and high z-score	(4) Intermediate z-score	(5) Low and high z-score	(6) Intermediate z-score	(7) Low and high KZ-index	(8) Intermediate KZ-index	(9) Low and high KZ-index	(10) Intermediate KZ-index
Analyst forecast $dispersion_{it-1}$*IDD_{it-1}	-0.460 * *		-0.587 *	0.444			-0.571 *	0.367		
	(−2.021)		(−1.715)	(0.834)			(−1.781)	(0.658)		
Analyst forecast $dispersion_{it-1}$	-0.292		-0.079	-0.790 * *			0.148	-0.574		
	(−1.105)		(−0.288)	(−2.027)			(0.472)	(−1.577)		
Analyst forecast $error_{it-1}$*IDD_{it-1}		-0.011 *			-0.020 * *	0.005			-0.024 * **	0.007
		(−1.688)			(−2.468)	(0.180)			(−3.204)	(0.278)
Analyst forecast $error_{it-1}$		0.001			0.001	-0.014			0.002	-0.029
		(0.215)			(0.078)	(−0.401)			(0.210)	(−1.163)
IDD_{it-1}	-0.013	-0.012	0.000	-0.034 * **	0.001	-0.033 * **	-0.006	-0.027 * **	-0.004	-0.028 * **
	(−1.412)	(−1.555)	(0.017)	(−3.056)	(0.049)	(−3.015)	(−0.307)	(−2.809)	(−0.215)	(−2.847)
Controls	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
Firm FE	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
Ind*Year FE	Y	Y	Y	Y	Y	Y	Y	Y	Y	Y
Observations	35,948	41,535	16,858	16,805	16,641	16,634	15,992	16,002	15,771	15,855
Adjusted R-squared	0.463	0.470	0.517	0.453	0.516	0.452	0.494	0.504	0.493	0.503

Note: This table reports the coefficient estimates of the interaction term between the *IDD* indicator variable and two proxies for information asymmetry: analyst forecast dispersion and analyst forecast error. Columns 1 and 2 utilize the full sample, while columns 3–10 use subsamples. The full sample is split into quartiles based on two default risk measures, Altman's (1968) z-score and KZ index (Kaplan and Zingales, 1997). Columns 3, 5, 7, and 9 report the results for the subsample consisting of the 1st and 4th quartiles and columns 4, 6, 8, and 10 the results for the subsample comprises the 2nd and 3rd quartiles. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * ** , * *, and * represent significance at the 1st, 5th, and 10th percentile levels, respectively.

6. Additional tests

In this section we explore how firms' characteristics influence the relation between labor mobility restrictions and short-term debt financing.

6.1. Innovation intensity

Firms in technology-intensive industries are more exposed to the threat of trade secret misappropriation, proprietary information protection to prevent losses due to information leakage is therefore more important for high innovation-intensive than for low innovation-intensive firms. We consequently argue that increased labor mobility restrictions resulting from adoption of the IDD should have a greater impact on short-term debt financing for the former firms. To test this conjecture, we employ two measures of innovation intensity: the ratio of R&D expenses to total assets and an indicator that equals one if a firm operates in high-tech industries following the definitions in Laitinen (2002), Clem et al. (2004), Kile and Phillips (2009), and Andrei et al. (2019). First, we partition our data into two subsamples based on the ratio of R&D expenses to total assets and examine the effect of the IDD for each subsample. As can be seen in columns 1 and 2 of Panel A of Table 13, the negative effect of the IDD adoption is concentrated in firms with high R&D expenses to total assets ratio. Next, we partition our data into subsamples based on 'high-tech' industries indicator variable. Columns 3 and 4 of Panel A of Table 13 show the negative effect of IDD adoption to present in the 'high-tech' firms subsample only. Overall, these results support our conjecture.

6.2. Risk of losing employees

Klasa et al. (2018) find that the impact of the IDD adoption on debt ratios is larger for firms with a greater ex-ante risk of losing employees who know their trade secrets to rivals. This risk is greater when firms face geographically closer rivals who can more easily poach their workers. We employ as the proxies for the ex-ante risk of a firm losing employees to rivals the weighted average distance between the firm's headquarters and each of its three-digit SIC industry rivals' headquarters (Klasa et al., 2018) and the percentage of the firm's industry rivals which are headquartered in the same state. We create two subsamples using these two proxies. Firms in the below (above) median weighted average distance subsample face geographically closer (farther) rivals, which indicates a greater (lower) ex-ante risk of losing employees to rivals.

Table 12
Agency costs mitigation mechanism.

Variables	(1) LT5	(2) LT5
Product market $fluidity_{it-1}$ *IDD_{it-1}	0.000	
	(0.310)	
Product market $fluidity_{it-1}$	0.002	
	(1.436)	
HHI_{it-1} *IDD_{it-1}		0.003
		(0.146)
HHI_{it-1}		-0.014
		(−0.986)
IDD_{it-1}	-0.011	-0.008
	(−0.997)	(−1.175)
Controls	Y	Y
Firm FE	Y	Y
Ind*Year FE	Y	Y
Observations	35,405	38,745
Adjusted R-squared	0.486	0.487

Note: This table reports the results of debt maturity regressions augmented with the proxies for agency costs. Columns 1 and 2 show the coefficient estimates of the interaction term with the two proxies for product market competitive threats (i.e., product market fluidity and HHI concentration metrics). Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * ** , * *, and * represent significance at the 1st, 5th, and 10th percentile levels, respectively.

While firms in the above (below) median percentage of rivals in the same state have more (fewer) rivals in the same state, indicating a greater (lower) ex-ante risk of losing employees to rivals. Columns 1–4 of Panel B of Table 13 show the coefficient of IDD_{it-1} to be negative and statistically significant at the 10% level for firms with geographically closer rivals or more rivals in the same state, which supports our conjecture.

6.3. Financial strength of rivals and asset specificity

The adoption of the IDD has a larger effect on leverage for firms that face financially stronger rivals and heightened competitive threats due to higher asset specificity (Klasa et al., 2018). Since firms that face financially stronger rivals are more likely to be exposed to the risk of losing key employees, we conjecture that the IDD adoption has a greater impact on debt maturity for these firms. We use the percentage of a firm's industry rivals with an S&P credit rating to test this conjecture. This is because unrated firms are financially weaker due to their limited or no access to bond markets (Harford and Uysal, 2014; Klasa et al., 2018). Supporting our conjecture, results in columns 1 and 2 of Panel C of Table 13 show that the impact of the IDD adoption on debt maturity to be concentrated in firms with higher percentage of rated rivals.

Furthermore, firms using more specific assets in their operation tend to face greater difficulty when selling these assets to raise funds to meet unforeseen financing needs (Klasa et al., 2018). As a result, asset specificity acts as an exit barrier for less productive firms, which results in excess capacity and increased product market competition (Klasa et al., 2018). We measure asset specificity using the three-digit SIC industry median ratio of machinery and equipment to book assets (Valta, 2012; Klasa et al., 2018), and partition our data into two subsamples. The results in columns 3 and 4 of Panel C of Table 13 support our conjecture that the IDD adoption has a stronger impact on debt maturity for firms with more specific assets.

7. New debt issues

Thus far we establish the negative causal relation between labor mobility restrictions and debt maturity using the maturity of all debts on a firm's balance sheet (i.e., balance sheet approach). In this approach, the debt maturity is the aggregation of the historical debt issuances (Guedes and Opler, 1996; Custodio et al., 2013). Guedes and Opler (1996) argue that the incremental approach relying on the debt maturity of new debt issues enables the identification of the determinants of debt maturity at all points along the maturity spectrum. To better investigate the changes in debt maturity structure caused by the IDD adoption that occurs over time, we follow Guedes and Opler (1996) by examining the impact of IDD adoption on the maturity of new debt issues.

We obtain new bond issues from Mergent FISD and private debt (i.e.,

Table 13
Additional tests.

Panel A: Innovation intensity

	R&D expenses		High-tech industry	
Variables	Below median	Above median	Non high-tech	High-tech
IDD_{it-1}	-0.011	-0.015 *	-0.008	-0.016 *
	(−1.490)	(−1.857)	(−1.403)	(−1.679)
Controls	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Ind*Year FE	Yes	Yes	Yes	Yes
Observations	38,972	35,530	46,725	28,400
Adjusted R-squared	0.506	0.500	0.494	0.468

Panel B: Ex-ante risk of losing employees to rivals

	Weighted average distance to rivals		**% rivals in the same state**	
Variables	Below median	Above median	Below median	Above median
IDD_{it-1}	-0.014 *	-0.012	-0.011	-0.014 *
	(−1.735)	(−1.378)	(−1.224)	(−1.915)
Controls	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Ind*Year FE	Yes	Yes	Yes	Yes
Observations	26,564	26,475	34,105	28,499
Adjusted R-squared	0.502	0.489	0.508	0.520

Panel C: Financial strength of rivals and asset specificity

	% Rated rivals		**Asset specificity**	
Variables	Below median	Above median	Below median	Above median
IDD_{it-1}	-0.004	-0.017 * *	-0.005	-0.017 *
	(−0.602)	(−2.590)	(−0.582)	(−1.958)
Controls	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Ind*Year FE	Yes	Yes	Yes	Yes
Observations	36,980	36,955	37,679	37,245
Adjusted R-squared	0.535	0.488	0.482	0.510

This table presents the results of additional tests using innovation intensity, ex-ante risk of losing employees to rivals, financial strength of rivals, and asset specificity. Panel A reports the results for innovation intensity subsamples. As the proxies for innovation intensity, we use the ratio of the R&D expenses to total assets and an indicator variable that equals one if a firm operates in the 'high-tech' industries. In Panel B, we define the weighted average distance to rivals as the asset-based weighted average of the distance (in kilometers) between a firm's headquarters and each of its three-digit SIC industry rivals' headquarters. The percentage of rivals in the same state is used as the second proxy. Panel C reports the results for the effect of the financial strength of rivals and asset specificity. We use the percentage of rivals with a credit rating as the proxy for the financial strength of rivals. The asset specificity is defined as the median ratio of machinery and equipment to total assets in a three-digit SIC industry across all years. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. ***, **, * represent significance at the 1st, 5th, and 10th percentile levels, respectively.

syndicated loans) from Dealscan database. The Mergent FISD sample contains 10,495 bond issues from 1987 unique firms, and their average (median) maturity is 11.99 (10) years. The Dealscan sample contains 29,816 loan facilities from 4,132 unique firms, and the average (median) maturity is 4.03 (5) years, which is much shorter than that of bond issues. Using these two sets of data, we re-run the DiD regression using the natural logarithm of maturity (in years) as the dependent variable. Following Custodio et al. (2013), we include issue type dummies or loan type dummies in all regressions and four decade indicator variables, such as *1980–1989 year indicator* which equals 1 if the bond was issued between 1980 and 1989 and zero otherwise, and the firm fixed effect.[10] The inclusion of the decade indicator variables enables the examination of the changes in intercepts over time (Custodio et al., 2013).

Column 1 of Table 14 reports the results for corporate bond issues and column 2 the results for syndicated loans. The coefficient of IDD_{it-1} is negative and statistically significant at the 1% level for corporate bonds, but not statistically significant for syndicated loans. These results indicate that firms tend to gravitate towards short-term debt in corporate bond markets following adoption of the IDD. This finding provides further confirmation of our hypothesis regarding the positive effect of limiting labor mobility on short-term debt financing.

Table 14
Initial maturity of new debt issues.

Variables	(1) Bond issues	(2) Syndicated loans
IDD_{it-1}	-0.052 * **	-0.005
	(−3.143)	(−0.310)
1980–1989 year indicator	-0.460 * **	
	(−5.874)	
1990–1999 year indicator	-0.553 * **	-0.034
	(−7.596)	(−1.284)
2000–2009 year indicator	-0.902 * **	-0.072 * *
	(−11.694)	(−2.224)
2010–2015 year indicator	-0.885 * **	0.008
	(−10.531)	(0.214)
Controls	Y	Y
Issue type dummies	Y	Y
Firm FE	Y	Y
Observations	10,495	29,816
Adjusted R-squared	0.307	0.645

Note: This table reports the results from regressing initial maturity of new bond issues and syndicated loans on the IDD changes. Column 1 (2) reports the effects of the IDD changes on the logarithm of the initial maturity (in years) of new bond issues (syndicated loans). Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * ** and * * represent significance at the 1st and 5th percentile levels, respectively.

8. Debt maturity dispersion

Our results support the hypothesis that firms prefer short-term debt to long-term debt following an increase in labor mobility restrictions. However, this raises the question whether increased labor mobility restrictions result in the concentration of maturity profile. To answer this question, in this section we examine the effect of labor mobility restrictions on debt maturity dispersion. When deciding the optimal debt maturity profile, firms tend to balance rollover risk against debt issuance cost and secondary market illiquidity (Choi et al., 2018). High issuance cost and illiquidity motivate firms to concentrate on a few large debt issuances, while rollover risk arising from the possibility of not being able to refinance large debts incentivizes firms to adopt a more disperse debt maturity profile. Using the credit rating downgrade of General Motors and Ford Motor Co. in 2005 as the quasi-natural experiment to gauge firms' response to the perception of high rollover risk, Choi et al. (2018) show that debt maturity dispersion of treated firms increases following the credit rating downgrade and the treated firms with high leverage experience substantial increase in debt maturity dispersion as compared to those with low leverage.

Similar to Choi et al. (2018) and Chiu et al. (2021), we obtain all types of corporate debt data from Standard & Poor's Capital IQ for the period 1989 – 2015. In line with Choi et al. (2018) and Chiu et al. (2021), we measure the degree of debt maturity dispersion as the inverse of the maturity profile's Herfindahl index. We follow the procedure in Chiu et al. (2021) to classify debt maturities into the nearest integer years using 13 buckets. For debts with maturity shorter than 10 years, we group them into 10 one-year maturity buckets (i.e., buckets 1–10). That is, debt maturities less than one year are grouped into bucket 1, debt maturities greater than one year but less than two years are grouped into bucket 2, and so on. Next, we group debt maturities from 11 to 15 years into bucket 11, debt maturities from 16 to 20 years into bucket 12, and the remaining debt maturities longer than 20 years into bucket 13.

Next, we define w_k $(w_{i,t,k} = x_{i,t,k}/\sum_{k=1}^{13} x_{i,t,k})$ as the proportion of principal amount maturing in each maturity bucket k, where x_k is the principal amount of debt for firm i maturing in each maturity bucket k. Next, we calculate firm i's Herfindahl index of debt maturity structure in year t as $HERE_{i,t} = \sum_{k=1}^{13} w_{i,t,k}^2$. Lastly, we calculate the maturity dispersion measure of firm i in year t using $Dispersion_{i,t} = 1/HERE_{i,t}$. A higher value of *Dispersion* indicates a more heterogenous maturity structure.

The average debt maturity dispersion for our sample is 1.37. We repeat the baseline regression by replacing the dependent variable with *Dispersion*. Column 1 of Table 15 shows that the estimated coefficient of IDD is negative and statistically significant at 5% level. This result indicates that the IDD adoption leads to a reduction in the dispersion of debt maturity. To examine whether the response to the IDD adoption is particularly strong in high leverage firms, we divide the full sample into two subsamples based on median leverage. We re-run the regression using subsamples of high and low leverage firms. Columns 2 and 3 of Table 15 show that the estimated coefficient of debt maturity dispersion is negatively significant for high leverage firms, but not for low leverage firms. These results indicate that high leverage firms indeed respond more strongly to increase in labor mobility restrictions following the IDD adoption by reducing their debt maturity dispersion.

9. Conclusion

In this paper, we examine the impact of labor mobility restrictions on firms' debt maturity. Employing the staggered-by-state adoption of the IDD as external shocks to labor mobility restrictions, we hypothesize that increased labor mobility restrictions can increase reliance on short-term debt financing by lowering default risk, aggravating information asymmetry, or elevating demand for external monitoring consequent to a reduction in product market threats, or reduce reliance on short-term debt to avoid rollover risk since less mobile managers are more risk averse. We find that firms headquartered in IDD-adopting states tend to hold more short-term debt post-IDD adoption. The positive causal effect of labor mobility restrictions on short-term debt financing stands up to a series of robustness tests.

Exploring the motives that might be at work by examining the default risk, information asymmetry, and agency costs mitigation mechanisms, we find evidence to support information asymmetry as the underlying mechanism that drives the negative relationship between labor mobility restrictions and debt maturity. Firms with high information asymmetry are more inclined to use short-term debt following adoption of the IDD. Default risk by itself does not seem to explain why firms hold more short-term debt post-IDD adoption. However, when we

[10] Issue types include enhancement, convertible, medium term note, asset-backed, Yankee, Canadian, foreign, Rule_144a, covenant, redeemable, putable, private placement, perpetual, and Rule_415. Loan types include term loans, revolver, 364-day facilities, and all other types included in DealScan database.

Table 15
Debt maturity dispersion.

Variables	(1) Full sample	(2) High leverage	(3) Low leverage
IDD_{it-1}	-0.061 * *	-0.095 * *	0.010
	(−2.406)	(−2.259)	(0.248)
Controls	Y	Y	Y
Firm FE	Y	Y	Y
Ind*Year FE	Y	Y	Y
Observations	19,732	10,354	8,227
Adjusted R-squared	0.382	0.415	0.369

Note: This table reports the results of the impact of IDD changes on debt maturity dispersion. The dependent variable is debt maturity dispersion. Column 1 uses the full sample and columns 2 and 3 use the subsample of firms with high and low leverage, respectively. Standard errors are clustered by the states in which firms are headquartered and t-values are shown in brackets. * * represents significance at the 5th percentile level.

consider information asymmetry and default risk in a single framework, we find some evidence of the non-monotonic relationship between default risk and debt maturity that supports Diamond's (1991) findings. We also do not find evidence consistent with using short-term debt as the governance mechanism to mitigate agency costs post-IDD adoption.

We further explore how firms' characteristics affect the relation between adoption of the IDD and short-term debt financing. Our results are more pronounced for firms with high innovation intensity, having a greater ex-ante risk of losing employees to rivals, facing financially stronger rivals, and facing heightened competitive threats due to higher asset specificity. Moreover, we consider the effect of labor mobility restrictions on new debt issues, and find the decline in long-term debt financing to mainly happen in corporate bond markets, but not in private debt markets. Finally, we explore if an increase in labor mobility restrictions incentivizes firms to concentrate on large debt issuance. Our results show that firms reduce their debt maturity dispersion post-IDD adoption. Our study thus complements the existing literature on labor mobility restrictions and debt maturity by showing how labor mobility restrictions affect corporate debt structure decisions.

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Data availability

The authors do not have permission to share data.

Appendix A. Variable definitions

Variables	Definition
LT1	Ratio of long-term debt to total debt.
LT2	Ratio of long-term debt minus debt maturing in 2 years to total debt.
LT3	Ratio of long-term debt minus debt maturing in 2 and 3 years to total debt.
LT4	Ratio of long-term debt minus debt maturing in 2, 3, and 4 years to total debt.
LT5	Ratio of long-term debt minus debt maturing in 2, 3, 4, and 5 years to total debt.
IDD	An indicator variable that equals one if firm *i* is headquartered in a state that has adopted the IDD from year *t-1*, and zero otherwise.
Tobin's Q	The sum of market value of equity plus book value of debt divided by total assets.
Leverage	Ratio of the sum of long-term debt and debt in current liabilities to total assets.
Size	The natural logarithm of total assets.
Profitability	Ratio of earnings before interest, taxes, depreciation, and amortization to total assets.
Credit rating	An indicator variable that equals one if the firm has an S&P long-term rating, and zero otherwise.
Abnormal earnings	Ratio of the difference between the income before extraordinary items, adjusted for common or ordinary stock (capital) equivalents for time t and t-1 to the market value of equity.
Tangibility	Ratio of gross property, plant, and equipment to total assets.
Asset volatility	Standard deviation of stock return during the fiscal year times market value of equity divided by market value of assets.
Asset maturity	Ratio of property, plant, and equipment to the sum of depreciation and amortization times the proportion of property, plant, and equipment in total assets, plus the ratio of current assets to the cost of goods sold times the proportion of current assets in total assets.
State GDP growth	The annual state GDP growth rate.
Political balance	The fraction of Congress members representing a given state in the U.S. House of Representatives that belong to the Democratic Party.
Analyst forecast dispersion	Standard deviation of earnings forecast scaled by the previous fiscal-year-end stock price.
Analyst forecast error	Absolute value of median forecast errors scaled by the previous fiscal-year-end stock price.
Product market fluidity	The measure of competitive threats faced by a firm in its product market.
HHI	The concentration measure that is positively associated with pricing power.
Z-score	Altman's (1968) z-score.
KZ index	-1.001909 * (IB+DP)/lagged PPENT + 0.2826389 * (AT+PRCC_F*CSHO–CEQ-TXDB)/AT + 3.139193 * (DLTT+DLC)/(DLTT+DLC+SEQ) – 39.3678 * (DVC+DVP)/lagged PPENT – 1.314759 * (CHE/lagged PPENT)
RtW	An indicator variable that equals one for states that have passed Right-to-Work legislation, and zero otherwise.
UTSA	An indicator variable that equals one if a firm is headquartered in a state in which UTSA is enacted, and zero otherwise.
R&D expenses	Ratio of the R&D expenses to total assets.
High-tech industry	An indicator variable that equals one if a firm is in the 'high-tech' industries, and zero otherwise.
Weighted average distance to rivals	The asset-based weighted average of the distance (in kilometers) between a firm's headquarters and each of its three-digit SIC industry rivals' headquarters.
% Rivals in the same state	The percentage of a firm's three-digit SIC industry rivals in the same state.
% Rated rivals	The percentage of a firm's three-digit SIC industry rivals with an S&P credit rating.
Asset specificity	The median ratio of machinery and equipment to total assets in a three-digit SIC industry.
Debt maturity dispersion	The inverse of the maturity profile's Herfindahl index following Choi et al. (2018) and Chiu et al. (2021).

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