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Alternative Work Arrangements and Cost of Equity: Evidence from a Quasi-Natural Experiment

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Abstract

I examine whether firms' use of alternative work arrangements, particularly temporary agency workers, affects their cost of equity. Exploiting a major labor-market deregulation in Japan that induced manufacturing firms to increase their employment of temporary agency workers, I show that the cost of equity decreased in manufacturing firms, relative to nonmanufacturing firms, after the deregulation. Further analysis using variations within manufacturing firms provides corroborating evidence. The rigidity in labor expenses and the cost of debt also decreased in manufacturing firms. Overall, alternative work arrangements increase the flexibility in labor costs, leading to lower operating leverage and cost of capital.

I. Introduction

Alternative work arrangements, or nonregular workers, such as temporary agency workers or fixed-term contract workers, have become increasingly prevalent in many countries. The *OECD Employment Outlook* (Organization for Economic Cooperation and Development (OECD (2014), Figure 4.1, p. 150)) shows that the incidence of nonregular employment increased in many countries over the 5-year period around the 2008 financial crisis. Katz and Krueger (2016) also document a sharp increase in alternative work arrangements in the U.S. labor

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market in recent years, showing that the percentage of workers in alternative work arrangements increased from 10.7% to 15.8% in the period of 2005–2015.[1]

Although firms are increasingly relying on alternative work arrangements, or nonregular workers, little is known about the potential effects of nonregular employment on the financial aspects of firms. One reason that firms utilize alternative work arrangements is the flexibility in the termination or continuation of the employment relationships with nonregular workers (Abraham and Taylor (1996), Segal and Sullivan (1997), Houseman (2001), Autor (2003), Ono and Sullivan (2013), and Cappelli and Keller (2013)). In general, employers have the flexibility to terminate the employment contract of a nonregular worker at minimal cost at the end of a contract period, which is not possible for a regular worker with an open-ended contract. Employers who plan to dismiss regular workers generally face high legal hurdles and/or are required to make high severance payments. The *OECD Employment Outlook* (OECD (2014), Table 4.4, pp. 165–167) documents the termination rules for regular and nonregular types of employment and shows that the termination rules for nonregular workers are far less strict than those for regular workers.

If alternative work arrangements help increase firms' flexibility in labor adjustment, firms' greater use of alternative work arrangements, or nonregular workers, would enable their cost structure to be more flexible and variable with respect to changes in demand for products. This increase in flexibility in the cost structure would effectively imply a decrease in operating leverage, which could affect the firms' cost of equity in at least two ways. On the one hand, a rise in alternative work arrangements, and an associated decrease in operating leverage, could cause a decline in the firms' existing exposure to systematic risk, potentially leading to a decrease in the cost of equity (e.g., Mandelker and Rhee (1984), Chen, Kacperczyk, and Ortiz-Molina (2011), Donangelo, Gourio, Kehrig, and Palacios (2019), and Favilukis and Lin (2016a), (2016b), among others). On the other hand, a rise in alternative work arrangements, or a decrease in the labor-induced operating leverage, may reduce the likelihood and the expected cost of firms' financial distress, which could encourage firms to increase their financial leverage to take advantage of the benefits associated with debt financing (Kuzmina (2018), Simintzi, Vig, and Volpin (2015), and Serfling (2016)). This implies that a rise in alternative work arrangements could lead to an increase in the cost of equity through an increase in the firms' financial leverage and equity beta. Thus, the overall effect of a rise in alternative work arrangements on the firms' cost of equity is not clear from these theoretical perspectives, and I view it as an empirical question. Specifically, I empirically investigate the *causal* effects of alternative work arrangements on firms' cost of equity.

An obvious challenge in estimating causal effects is the endogeneity between firms' cost of capital and their employment decisions, which are likely to be jointly determined. To circumvent the endogeneity problem, I exploit a unique feature of a plausibly exogenous labor-market deregulation in Japan in 2003. The 2003 amendment to the Worker Dispatching Act in Japan is an ideal setting to

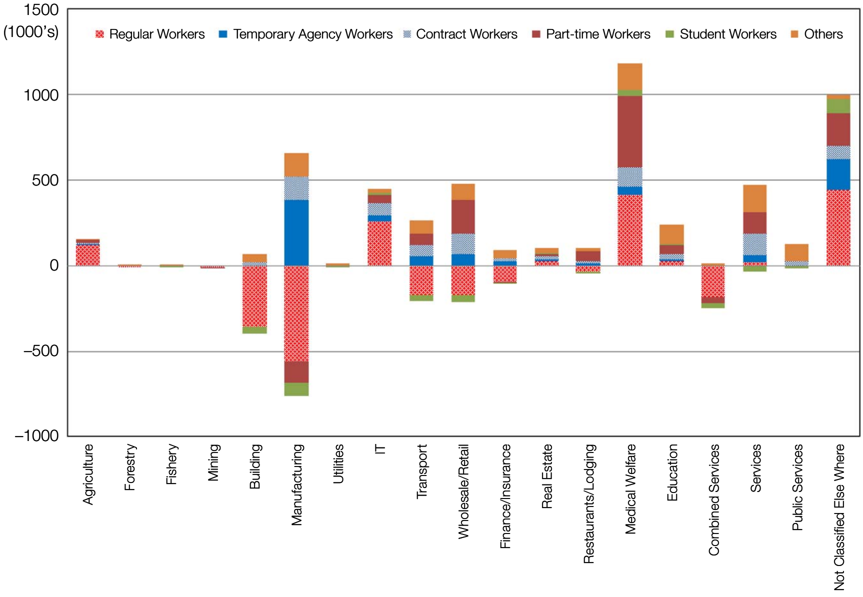
[1]See also Katz and Krueger (2019) for their updated findings on the recent trends in alternative work arrangements in the United States.

establish causality by clearly distinguishing between the treatment group and the control group. The act, enacted in 1986, specifies rules on the operations of temporary work agencies in Japan. Before the amendment in 2003, the act had prohibited temporary work agencies from dispatching their workers to production lines in manufacturing plants. However, given the stagnant macroeconomic and labor-market conditions in Japan in the early 2000s, the government lifted the ban in 2003, allowing temporary agency workers to engage in production line work in manufacturing.[2] This deregulation had a clear and large effect on the employment decisions of firms in the manufacturing sector but, importantly, not on those of firms in other sectors. After the deregulation, the number of temporary agency workers sharply increased for firms in the manufacturing sector but changed little in other sectors (see Figure 1).[3]

Using this major labor-market deregulation as a quasi-natural experiment, I infer the causal effects of alternative work arrangements on the firms' cost of equity in a difference-in-differences (DID) framework. Given the unique feature of

FIGURE 1
Changes in the Number of Workers by Types of Employment in Broad Industry Categories

Figure 1 shows the changes in the number of workers by types of employment between 2002 and 2007 in 19 broad industries. The types of employment are regular workers, temporary agency workers, contract workers, part-time workers, student workers, and others. I constructed the figure using official statistics from the 2002 and 2007 Employment Status Survey, which is published every 5 years by the Japan Statistics Bureau.



[2]I discuss the institutional background of the deregulation in detail in the next section.

[3]The data used to construct Figure 1 are from the government official statistics in the Employment Status Survey, published by the Japan Statistics Bureau every 5 years. Figure 1 shows a clear picture of the effects of the amendment on labor markets. Between 2002 and 2007, the number of temporary agency workers increased from 195,700 to 580,600 in firms in manufacturing industries but did not change significantly in other industries.

the amendment in 2003, which affected only manufacturing firms, I define firms in the manufacturing sector as the treatment firms and those in other sectors as the control firms. I adopt a standard DID approach in panel regressions, using a sample of public firms in Japan in the period of 2000–2006. I use the implied cost of equity of Gebhardt, Lee, and Swaminathan (2001) as the primary measure of the cost of equity. I then examine how the amendment in 2003 affected the cost of equity of firms in the treatment and control groups. Using a DID framework, I find that the cost of equity decreased in firms in manufacturing industries, relative to firms in other industries, after the deregulation in 2003. I obtain this result after including a host of other potential determinants of the cost of equity, firm fixed effects, and year fixed effects to mitigate potential biases arising from omitted variables. Note that firm fixed effects will absorb any time-invariant differences between manufacturing and nonmanufacturing firms. An analysis using a matched sample between manufacturing and nonmanufacturing firms by an entropy-balancing method also yields a similar result. This finding is most consistent with the interpretation that an (expected) increase in flexibility in labor costs and a decrease in operating leverage through temporary agency workers in manufacturing firms after the deregulation caused a decrease in the firms' exposure to systematic risk and the cost of equity at the time of the deregulation in 2003. My most preferred specification suggests that the cost of equity decreased by 1.085% in manufacturing firms after the deregulation.

I recognize that my causal interpretation of the finding crucially hinges on the validity of the parallel-trends assumption in my setting. A potential concern, however, is comparability between the treatment-group firms (i.e., manufacturing firms) and the control-group firms (i.e., nonmanufacturing firms). For instance, it is possible that some omitted or unobservable differences in time-varying characteristics between the two groups may have driven a difference in the changes in the cost of equity, even in the absence of the deregulation. These time-varying differences might also have endogenously induced the government to adopt a deregulation specifically targeting the treatment-group firms.

I address these concerns in several ways. First, I control for their time-varying differences in sales and earnings at the industry level as well as differences in past stock returns and exposures to foreign exchange fluctuations between the two groups in the estimations. Second, I check whether the pretreatment trends of the cost of equity in the treatment-group firms and the control-group firms were similar before the deregulation because the parallel-trends assumption would not likely be valid if their trends in the cost of equity were different before the deregulation. Third, I exploit cross-sectional variations, only *within* manufacturing firms, to test the effects of the deregulation on the cost of equity. Specifically, I examine whether the effects of the deregulation on the cost of equity are more pronounced for manufacturing firms with greater exposure to the deregulation. Firms that produce goods primarily in manufacturing plants located in Japan and those that employ labor input intensively in production should have been strongly affected by the deregulation targeted at temporary agency workers residing in Japan. These variations within manufacturing firms also enable me to control for any unobservable time-varying differences in the characteristics of manufacturing firms and nonmanufacturing firms. The results from these additional tests confirm that omitted or unobservable time-varying differences between manufacturing and nonmanufacturing firms are not responsible for my main results.

I also conduct a number of analyses to cement the validity of my interpretation that the cost of equity decreased in manufacturing firms due to the deregulation related to temporary agency workers. Although the overall number of temporary agency workers increased in the manufacturing sector after the deregulation, the size of an increase in temporary agency workers varied across industries within the manufacturing sector. If my findings truly reflect the effects of deregulation on the cost of equity through temporary agency workers, the cost of equity should have decreased more in firms in industries in which the number of temporary agency workers increased more after the deregulation within the manufacturing sector. Indeed, I find evidence that is consistent with this conjecture.

Another approach to validating my interpretation is to examine the effects of the deregulation on the flexibility in labor costs. Because I argue that the cost of equity decreased due to an expected increase in labor-cost flexibility, I examine the effects of the deregulation on the labor-cost flexibility in manufacturing firms, relative to nonmanufacturing firms, in two ways. First, I examine whether the sensitivity of labor expenses to sales increased after the deregulation in manufacturing firms. An increase in flexibility in labor costs through temporary agency workers would enable manufacturing firms to adjust labor expenses to demand fluctuations more easily after the deregulation than before. Consistent with this conjecture, I find that labor expenses (firm profits) became more (less) sensitive to sales in manufacturing firms after the deregulation. Second, I examine whether the volatility of labor expenses increased in manufacturing firms after the deregulation. An increase in labor-cost flexibility could imply higher (lower) volatility in labor expenses (firm profits) in manufacturing firms after the deregulation. I find evidence consistent with this expectation. Overall, these results suggest a decrease in labor leverage in the firms affected by the deregulation (Donangelo et al. (2019), Favilukis and Lin (2016a), (2016b), and Favilukis, Lin, and Zhao (2020)).

Finally, I examine whether the deregulation affected the firms' cost of debt. Increased flexibility in labor costs could mitigate adverse cash-flow shocks in bad times, potentially leading to a firm's lower default risk and cost of debt. Indeed, after controlling for various issue- and firm-level characteristics, I find that the yields of corporate bonds at the time of issuance decreased in manufacturing firms after the deregulation.

Taken together, my results are most consistent with the interpretation that the deregulation in the labor market decreased the cost of capital in manufacturing firms by decreasing the operating leverage and exposure to systematic risk in those firms. My study contributes to the growing literature on labor and finance. A number of recent studies in corporate finance examine the causal effects of changes in labor laws or labor-related government policies on the corporate capital structure (Matsa (2010), Agrawal and Matsa (2013), Kuzmina (2018), Simintzi et al. (2015), and Serfling (2016)) or other dimensions (e.g., Acharya, Baghai, and Subramanian (2013) examine effects on innovation, and Dessaint, Golubov, and Volpin (2017) examine effects on mergers and acquisitions (M&As), among others). There is also a line of research investigating the effects of unionization on various aspects of corporate finance, including financing decisions (Klasa, Maxwell, and Ortiz-Molina (2009), Matsa (2010), and Schmalz (2015)), M&As (John, Knyazeva, and Knyazeva (2015)), and executive compensation (Huang, Jiang, Lie, and Que (2017)), among others. My study complements these studies but differs from them in that I focus on alternative work arrangements, particularly temporary agency workers, and investigate their causal effects on the cost of capital.

I also recognize that several recent articles on asset pricing examine the cross-sectional relationships between labor and stock returns, including the effects of labor-market frictions, adjustment costs, and search costs (Belo, Lin, and Bazdresch (2014), Belo, Lin, Li, and Zhao (2017), and Kuehn, Simutin, and Wang (2017)), wage rigidity and scale inflexibility (Donangelo et al. (2019), Favilukis and Lin (2016a), (2016b), and Gu, Hackbarth, and Johnson (2018)), and organization capital and labor mobility (Eisfeldt and Papanikolaou (2013), Donangelo (2014)). Some of these studies link labor to finance via the operating leverage induced by labor in a number of contexts. My article is also motivated by operating leverage originating from labor but in a unique context: flexibility in labor contracts with temporary agency workers. Furthermore, my article differs from most of these asset pricing articles on equilibrium relationships between labor and expected stock returns because my primary aim is to establish *causality* from labor to expected returns by exploiting a plausibly exogenous change in labor laws.

My article is most closely related to those by Chen et al. (2011), Donangelo (2014), Donangelo et al. (2019), and Favilukis and Lin (2016a), (2016b), which examine how a labor-induced form of operating leverage affects the cost of equity or expected stock returns.[4] Although I share conceptually similar interests in the effects of labor-induced operating leverage on the cost of financing, my article adopts a different empirical strategy and pays particular attention to estimating the *causal* effects of the labor-induced operating leverage on the cost of equity. The empirical measures of the labor-induced operating leverage employed in previous articles were primarily based on cross-sectional variations across industries, such as the industry-level unionization rate (Chen et al. (2011)), industry-level labor mobility (Donangelo (2014)), or industry-level labor share (Donangelo et al. (2019), Favilukis and Lin (2016a)). Although these cross-sectional variations across industries are useful for identification, there is a possibility that some omitted and/or unobservable industry-level characteristics, which are correlated with the industry-level measures, may bias the estimation results. My identification strategy, which uses a DID framework, adds to the existing empirical methods in the previous articles because I not only utilize cross-industry variations (between the treatment group and the control group) but also exploit a (time-series) plausibly *exogenous* shock, which is specific to labor, arising from a law change. I believe that my empirical design will further help establish causality from labor-induced operating leverage to expected stock returns.

As far as I can determine, few researchers directly link alternative work arrangements to finance. A notable exception is Kuzmina (2018), who infers a causal relationship between the fraction of a firm's fixed-term workers and

[4]Relatedly, Campello, Gao, Qiu, and Zhang (2018) examine the effects of union elections on bond prices. Alimov (2015) examines the effects of employment protection on loan spreads.

corporate capital structure. She exploits adoptions of government subsidy programs in Spain that encouraged the conversion from temporary contracts to permanent contracts and uses an instrumental variable (IV) estimation strategy with the expected subsidy as the instrument. My article and Kuzmina's (2018) share a similar interest in the effects of temporary workers on corporate finance, but the focus of my article is the cost of capital. Furthermore, my identification strategy is different because I exploit a change in a labor law that is directly related to temporary agency workers to conduct a DID analysis. Moreover, I use a large-scale deregulation that encouraged firms to increase temporary workers, rather than decrease them, in contrast to the government programs used by Kuzmina (2018).

Overall, my article contributes to the growing literature on labor and finance by investigating the *causal* effects of alternative work arrangements, particularly temporary agency workers, on the cost of capital using a major labor-market deregulation on temporary agency workers in Japan as a quasi-natural experiment. The remainder of the article is organized as follows: Section II describes the institutional details behind the deregulation, develops a conceptual framework, and discusses the relationship of my article to the existing literature. Section III describes data and constructs the variables. Section IV describes the empirical analyses and discusses the results. Section V concludes.

II. Institutional Background and Conceptual Framework

A. Amendment to the Worker Dispatching Act in 2003

In general, it is quite difficult to terminate the contracts of, or dismiss, regular workers in Japan. Regular (or permanent) workers have an open-ended contract with an employer that can be terminated only under very restricted circumstances. Sugeno and Yamakoshi (2014) describe the historical development of Japanese labor laws, including the Labor Standard Act enacted in 1947 and the related case law, which tend to strongly protect the employment of regular workers. Suppose, for instance, that a firm would like to conduct a restructuring of its business operations that could involve the dismissal of employees through downsizing or shutting down some divisions and operations to improve operational efficiency and profitability. In this case, corporate restructuring is unlikely to be a valid reason for the dismissal of employees if there is no corroborative evidence of a firm's sufficiently poor financial and operational performance.[5]

Given the difficulty of dismissing regular workers in Japan, firms have been increasingly replacing regular workers with nonregular workers, such as part-time workers, temporary agency workers, and fixed-term contract workers. The fraction of nonregular workers over total workers in Japan increased from 19.7% to 35.5% in the period of 1987–2007, according to government statistics. In general, firms incur minimal costs when terminating an employment relationship with a nonregular worker at the end of a contract period in Japan. In terms of temporary

[5]One may wonder whether firms could alternatively scale down the wages and/or working hours of regular workers, which are normally specified in employment contracts. However, in light of the Labor Contracts Act and the related case law, a reduction of the wages and/or working hours of regular workers, which negatively affects their labor incomes, will be very difficult in practice without consent from the regular workers.

agency workers, the Worker Dispatching Act, established in 1985, governs the operations of temporary work agencies and the working conditions of their workers. Initially, the act allowed only 13 occupations on the "positive" list for temporary work agencies to dispatch their workers to workplaces. These occupations include some types of administrative workers, secretaries, translators, and software developers. The agencies were not allowed to send their workers for occupations that were not on the list.

Although the list initially included only 13 occupations, the government increased the number of occupations on the positive list to 26 in 1996.[6] Occupations included in the initial list at the establishment of the act in 1985 and those in the expanded list in 1996 were more or less relevant to the firms' employment of temporary agency workers in all industries.[7] For example, even though production line work in manufacturing was not on the list, manufacturing firms were free to hire temporary agency workers for any other occupations, such as administrative workers. In 1999, the list was abolished, and with a few exceptions, it became legal for the agencies to assign their workers to most occupations. This is equivalent to expanding the list to include almost any occupation other than the exceptions. Those exceptions were provided on a new list, called the "negative" list, created by the government. These exceptions, to which the agencies were prohibited from assigning their workers, included production line work in manufacturing, construction work, port and harbor transportation services, medical services, and security services.[8]

However, the amendment to the Worker Dispatching Act in 2003 lifted a ban that had prohibited temporary work agencies from dispatching workers to manufacturing lines since the enactment of the act in 1985.[9] Admittedly, the amendment to the Worker Dispatching Act did not come unexpectedly. The cabinet office of the government under the ruling party, the Liberal Democratic Party (LDP), set up the Council for Regulatory Reform in Apr. 2001. The main task of the council was to comprehensively examine the red tape in Japan, or excess bureaucracy and regulation, that generated additional business costs in all sectors, including the labor market. In particular, the second report from the council, which was issued in Dec. 2002, strongly recommended lifting the ban that had prohibited temporary agency workers from working at manufacturing plants. This recommendation from the council was considered a key development for the process of

[6]The added occupations in the expanded list in 1996 include sales workers, research and business assistants, copy editors, and designers, among others.

[7]This makes it difficult to find control firms that were not affected by the establishment of or the subsequent changes in the list; therefore, I cannot include these changes in the list before 2003 as natural experiments in my DID analysis.

[8]Although the specific reasons for the exclusion of those occupations from temporary agency workers varied across occupations, regarding production line work in manufacturing, an important consideration for its exclusion appeared to be requests from labor organizations in manufacturing sectors, which consisted predominantly of regular workers.

[9]Kuroki (2012) also provides a historical overview of employment practices in Japan and the amendment to the Worker Dispatching Act in 2003.

deregulation of temporary agency workers, ultimately leading to the formal passage of a related bill by the parliament in June 2003.[10]

The amendment went into effect in Mar. 2004, and it became legal for manufacturing firms to hire temporary agency workers as production line workers after Mar. 2004. The fractions of temporary agency workers over total workers were 1.4% in 2002 and 3.0% in 2007, indicating that temporary agency workers accounted for a relatively small portion of total workers.[11] However, the growth rate of temporary agency workers was striking. Although the number of total workers in all sectors increased by only 4.7% in 2002–2007, the number of temporary agency workers increased by 123.0% in the same period. Furthermore, focusing on the manufacturing sector, whereas the total number of workers actually *declined* by 2.3% in the manufacturing sector in the same period, the number of temporary agency workers in the same sector increased by 196.7%.[12]

B. Wage Rigidity and Contract Periods of Regular and Nonregular Workers in Japan

The *OECD Employment Outlook* (OECD (2014)) defines nonregular workers as those workers who do not benefit from the same degree of protection against contract termination that regular workers do. Regular workers normally have an open-ended employment contract that cannot be easily terminated by employers, whereas nonregular workers work under a fixed-term contract that can be terminated by employers at minimal cost at the end of a contract period. This difference is particularly significant in Japan, in which the bar for dismissing regular workers is set high, whereas there are fewer restrictions on terminating an employment contract with a nonregular worker at the end of a contract period. In general, the shorter contract period of nonregular workers would enable firms to make more frequent adjustments to the wages and employment of nonregular workers.

I posit that an increase in nonregular workers in manufacturing firms after the labor-market deregulation caused a subsequent decrease in the rigidity of those firms' labor expenses, operating leverage, and cost of equity. To validate my conceptual framework, it is important to show that the wages of nonregular workers are actually less rigid than those of regular workers in Japan. In this section, I examine and compare the rigidity in real wages between nonregular workers and regular workers. I also examine the typical contract periods of regular and nonregular workers in the labor markets in Japan.

1. Wage Rigidity

The analysis of wage rigidity requires historical data on the wages of regular and nonregular workers, which are publicly available from the Basic Survey on

[10]In my subsequent empirical analysis, I estimate the effects of the deregulation on a firm's change in the cost of equity between June 2002 (i.e., 6 months before the council's recommendations in Dec. 2002) and June 2003 (i.e., the month in which the amendment was approved by the parliament) to eliminate any potential effects of the council's recommendations in Dec. 2002 on the cost of equity.

[11]In manufacturing, the fractions of temporary agency workers were 1.9% in 2002 and 5.7% in 2007.

[12]This increase in the number of temporary agency workers by 196.7% in the manufacturing sector is almost twice as large as the contemporaneous increase by 95.5% in nonmanufacturing sectors during 2002–2007.

Wage Structure published annually by the Ministry of Health, Labor, and Welfare (MHLW). The survey collects information on the payroll of employees from a sample of business establishments in the month of June every year and reports the average monthly wages per employee in June, as well as other payroll information, every year at the broad industry level. The MHLW started the survey in 1948 but did not report separate figures for regular and nonregular workers until 2005. Thus, I collect data on the wages of both types of workers from 2005 to 2018 (i.e., the latest available year). Then, I use the data to compute the annual growth rates of the average monthly real wages per employee from a previous year for 12 broad industries between 2006 and 2018.[13] To measure the wage rigidity, following Favilukis and Lin (2016a), I compute the standard deviations of the annual growth rates of the average monthly real wages per employee during the sample period.

Table 1 reports the (time-series) means and standard deviations of the annual growth rates of the average monthly real wages per employee for regular and nonregular workers for 12 broad industries during the sample period (2006–2018). If the wages of nonregular workers are less rigid than the wages of regular workers, the standard deviations of the wage growth of nonregular workers should be higher than those of regular workers. The results in Table 1 are generally consistent with this expectation. In most of the industries, the standard deviations of the real wage growth of nonregular workers are higher than those of regular workers. The last row reports the Levene's robust test statistic for the equality of variances of real wage growth between regular and nonregular workers. The p-value of the test statistic is almost 0, indicating that the variance of the real wage growth of

TABLE 1

Real Wage Rigidity of Regular and Nonregular Workers by Industry

Table 1 reports the (time-series) means and standard deviations of the annual growth rates of the average monthly real wages per employee for regular and nonregular workers for 12 broad industries during 2006–2018. The payroll data are from the Basic Survey on Wage Structure published by the Ministry of Health, Labor, and Welfare (MHLW).

Real Wage Growth ($=W_t/W_{t-1}$)	Regular Workers			Nonregular Workers		
	Mean	Std. Dev.	N	Mean	Std. Dev.	N
Real estate	0.993	0.026	13	0.997	0.032	13
Medical welfare	0.997	0.016	13	1.004	0.020	13
Wholesale/retail	0.997	0.015	13	1.006	0.029	13
Restaurants/lodging	0.996	0.007	13	1.003	0.013	13
Construction	0.999	0.023	13	0.999	0.030	13
Information technology	1.000	0.050	13	1.014	0.059	13
Education	0.994	0.025	13	1.007	0.035	13
Manufacturing	0.997	0.019	13	1.007	0.019	13
Transportation	0.999	0.029	13	1.005	0.027	13
Finance/insurance	0.996	0.034	13	1.001	0.083	13
Mining	1.002	0.034	13	1.036	0.067	13
Utilities	0.999	0.019	13	1.000	0.058	13
Total	0.997	0.026	156	1.007	0.044	156

Levene's test statistic (for equality of variances) = 19.675 (p-value = 0.000)

[13] I convert nominal wages to real wages using the Consumer Price Index (CPI) published annually by the Japan Statistics Bureau to adjust for changes in the general price level.

nonregular workers is significantly higher than that of regular workers.[14] Table A1 in the Supplementary Material reports the corresponding statistics with an alternative definition of average monthly real wages per employee, which adds the average monthly overtime pay and the average amount of annual bonus divided by 12, both in real terms, per employee to the average monthly real wages in Table 1. The results in Table A1 are similar to those in Table 1.

2. Contract Periods

Next, I turn my attention to the contract periods of regular and nonregular workers. The labor-market deregulation in 2003 in Japan removed the ban that had prohibited temporary agency workers from engaging in production line work in manufacturing firms. The deregulation effectively increased the firms' flexibility in their choice of contract periods for production line workers in manufacturing firms. A shorter contract period of nonregular workers, in general, would enable firms to make more frequent adjustments of both the wages and employment of nonregular workers. In this section, I examine the typical contract periods of regular and nonregular workers in Japan.

Regarding the contracts of regular workers in Japan, by definition, they have an open-ended contract that does not specify the length of the contract period. This effectively implies that, given very stringent termination rules concerning a firm's dismissal of regular workers, the contract would remain in effect until a regular worker voluntarily leaves a firm or retires at the firm's retirement age, which is typically the age of 60 in Japan. As an example, for a recent college graduate entering the labor force as a regular worker at the age of 22, a contract period for such a worker can be effectively regarded as 38 years, if the worker intends to stay in the same firm until retirement.[15]

In contrast to the long-term contracts of regular workers, nonregular workers tend to have short-term contracts. To provide more details, I show the frequency distributions of the contract periods of some types of nonregular workers in Japan. I obtain aggregate data on the contract periods of nonregular workers from the latest Employment Status Survey conducted by the Japan Statistics Bureau. Conducted every 5 years, this household survey collects various types of information about the employment status of household members. The survey reports the frequency distributions of the contract periods of nonregular workers. I present the distributions reported in the latest survey conducted in 2017. The contract periods are classified as follows: i) up to 1 month, ii) 1–3 months, iii) 3–6 months, iv) 6 months to 1 year, v) 1–3 years, vi) 3–5 years, and vii) more than 5 years.

Table 2 reports i) the overall frequency distribution of the contract periods of nonregular workers, ii) the distribution of temporary agency workers, and iii) the distribution of nonregular workers directly hired by firms (directly hired nonregular workers; e.g., contract workers and part-time workers). According to

[14] In unreported results, I also find that the standard deviation of the employment growth of nonregular workers is significantly higher than that of regular workers, suggesting lower rigidity of employment in nonregular workers than regular workers.

[15] Regarding the salaries of regular workers, they are normally subject to a firm's internal rules of employment. The salary level would depend on the title and responsibility of the worker's job within a firm and would increase as he or she moves up the corporate ladder.

TABLE 2
Contract Periods of Nonregular Workers

Table 2 reports the overall frequency distribution of the contract periods of i) nonregular workers, ii) temporary agency workers, and iii) nonregular workers directly hired by firms (directly hired nonregular workers, e.g., contract workers, part-time workers). The data are from the Employment Status Survey conducted by the Japan Statistics Bureau in 2017.

Contract Periods	All Nonregular Workers	Temporary Agency Workers	Directly Hired Nonregular Workers
≤ 1 month	0.83%	2.44%	0.67%
1–3 months	10.24%	46.79%	6.59%
3–6 months	18.19%	19.75%	18.03%
6–12 months	47.11%	14.19%	50.40%
1–3 years	18.59%	13.47%	19.10%
3–5 years	3.27%	2.49%	3.35%
> 5 years	1.76%	0.87%	1.85%

Table 2, a typical contract period of a nonregular worker is between 6 months and 1 year (47.11%, usually 1 year), followed by 1–3 years (18.59%) and 3–6 months (18.19%). Although the distribution of directly hired nonregular workers exhibits a similar pattern to the overall distribution, it is interesting to note that the most common contract period of temporary agency workers is 1–3 months (46.79%, usually 3 months). Temporary agency workers appear to give firms more flexibility in terms of contract length than other types of nonregular workers.

C. Conceptual Framework and Related Work

One reason that employers often cite for using temporary agency workers is the high flexibility in labor adjustment made possible by hiring such workers. The existing studies on temporary agency workers document that firms use flexible staff arrangements to accommodate demand fluctuations (Abraham and Taylor (1996), Segal and Sullivan (1997), Houseman (2001), Autor (2003), Ono and Sullivan (2013), and Cappelli and Keller (2013)). Because it is easier for firms to adjust their labor forces through nonregular workers than through regular workers, firms can manage their labor expenses flexibly by hiring nonregular workers, which can make the firms' labor costs variable with fluctuations in demand for their products. Thus, the use of nonregular workers reduces the rigidity associated with labor costs and increases the flexibility to adjust the labor force and expenses along with demand fluctuations. This increase in flexibility essentially implies a decrease in operating leverage, or the portion of fixed costs in total costs. Mandelker and Rhee (1984), among others, show that higher operating leverage makes firms' cash flows and returns to shareholders more variable, which amplifies their existing exposure to systematic risk and increases the cost of equity (e.g., Mandelker and Rhee (1984), Carlson, Fisher, and Giammarino (2004), Novy-Marx (2011), Chen et al. (2011), Eisfeldt and Papanikolaou (2013), Donangelo (2014), Favilukis and Lin (2016a), (2016b), Gu et al. (2018), and Donangelo et al. (2019)).

Specifically, several studies examine the link between the labor-induced operating leverage and the cost of equity or expected stock returns. Chen et al. (2011) examine the effects of labor unions on the cost of equity and argue that unions make wages sticky and layoffs costly, which increases firms' labor-induced operating leverage. They find that the cost of equity is higher for firms in industries with stronger labor unions. Relatedly, under certain assumptions on wage rigidity and capital–labor complementarity, Donangelo et al. (2019) and Favilukis and Lin (2016a), (2016b) argue that the labor share, or the ratio of labor expenses over value added, can be a valid proxy for capturing the labor-induced operating leverage. They show that firms (or industries) with a high labor share tend to have high expected stock returns. In a different context, Donangelo (2014) argues that labor mobility creates a labor-induced form of operating leverage and affects the expected stock returns. A higher (lower) labor mobility of workers across industries with general labor skills (industry-specific labor skills) would make the wages of those workers less (more) elastic to industry-specific shocks because those workers would be more (less) able to search for higher wages across many industries. Thus, firms in industries that depend mostly on workers with general labor skills would face less elastic wages of workers, which would make the level of cash flows to shareholders more sensitive to industry-specific shocks and amplify the firms' existing exposure to systematic risk. Overall, these studies predict a positive relationship between labor-induced operating leverage and the cost of equity.

If firms' use of nonregular workers, or temporary agency workers, decreases operating leverage, such a decrease in labor-induced operating leverage would directly imply a decrease in the cost of equity through a reduction in their existing exposure to systematic risk. However, some studies on labor and capital structure suggest that there may be an *indirect* effect of a decrease in the labor-induced operating leverage on the cost of equity, which could potentially imply an increase in the cost of equity. Based on the trade-off relationship between operating leverage and financial leverage as noted by Mandelker and Rhee (1984), recent studies on labor and finance examine whether a labor-induced form of operating leverage crowds out financial leverage (Kuzmina (2018), Simintzi et al. (2015), and Serfling (2016)). An increase in operating leverage may lead to an increase in the likelihood and the expected cost of firms' financial distress, which could reduce the firms' capacity to take on debt. Indeed, these articles generally find that an exogenous increase in labor-induced operating leverage leads to a decrease in financial leverage, consistent with the trade-off between operating and financial leverage. If an increase in firms' use of temporary agency workers, or a decrease in labor-induced operating leverage, leads to an increase in financial leverage, which could be driven by the firms' optimal debt policies to take advantage of the benefits associated with debt financing, the firms' cost of equity might also increase through the well-known theoretical link between financial leverage and equity beta.

Thus, the literature suggests at least two ways through which temporary agency workers could affect firms' cost of equity. On the one hand, a reduction in labor-induced operating leverage would, assuming no change in financial leverage, decrease the cost of equity by reducing the firms' existing exposure to systematic risk. I call this channel the *direct effect*. On the other hand, if a reduction in labor-induced operating leverage induces firms to increase financial leverage, the cost of equity could increase because the financial leverage would increase the equity beta. I call this channel the *indirect effect*. Thus, the net effect of an increase in temporary agency workers on the cost of equity is not clear, depending on which effect dominates the other, which is ultimately an empirical question. I attempt to answer this question by estimating the net effect of temporary agency workers, or alternative work arrangements more generally, on the cost of equity by exploiting a plausibly exogenous shock (i.e., a labor-market deregulation event) to the supply of those workers in the manufacturing sector.

III. Data

A. Sample Selection

I obtain historical financial data and stock prices from the Nikkei NEEDS FinancialQUEST 2.0 database, which is widely considered the most comprehensive financial database on public firms available in Japan. My base sample covers the period of 2000–2006. I include all public companies except those in the financial and utilities industries. The final sample is an unbalanced panel comprising 13,112 firm-year observations.

B. Variable Definitions

1. Cost of Equity

My primary measure of the cost of equity is the implied cost of equity by Gebhardt, Lee, and Swaminathan (2001) (GLS). The literature documents that realized returns, which are a backward-looking measure by construction, are a noisy proxy for expected returns. For my purpose, it is appropriate to use the implied cost of equity, which is a forward-looking measure, because the news about the passage of the deregulation bill in parliament would have been incorporated into the stock prices and expected future cash flows. Numerous accounting and finance studies use the implied cost of equity as a primary measure for expected stock returns.[16]

Computation of the implied cost of equity requires information on expected future earnings, and the literature typically uses analyst forecasts as estimates of firms' future cash flows. Although I largely follow a conventional methodology to calculate the implied cost of equity, the limited availability of analyst forecast data in my sample period (i.e., early 2000s) for Japanese firms requires me to consider a proxy for firms' expected earnings other than analyst forecasts.[17] Therefore, in my main analysis, I compute my primary measure of the cost of equity, GLS, by following an approach suggested by Hou, Van Dijk, and Zhang (2012), which proposes a cross-sectional model to generate earnings forecasts. The basic idea of the forecasting method of Hou et al. (2012) is to estimate a cross-sectional regression model of earnings on a number of lagged explanatory variables using the

[16]In addition to the GLS implied cost of equity, I also consider four alternative measures of the implied cost of equity by Claus and Thomas (2001) (CT), Ohlson and Juettner-Nauroth (2005) (OJN), Easton (2004) (modified price-earnings growth (MPEG)), and Gordon and Gordon (1997). The unreported results with these alternative measures of the implied cost of equity are qualitatively similar to my main results with the GLS implied cost of equity.

[17]Data on analyst forecasts are available only for approximately 30% of my sample firms in the early 2000s in Japan. The accounting literature also notes several issues concerning the predictive power or biases of the implied cost of equity, which potentially arise from the quality of the analyst forecasts.

previous 10 years of observations and use estimated coefficients to predict future earnings based on the current values of those explanatory variables.[18] I compute the GLS implied cost of equity for each firm-year using a firm's closing stock price on June 30 and the most recent accounting data available before June 30 in each year, which typically come from the financial statements reported on Mar. 31 every year (i.e., the most commonly used fiscal year-end date for Japanese firms). As a robustness check, I also compute the estimates of the GLS implied cost of equity based on Institutional Brokers' Estimate System (IBES) consensus analyst forecasts to examine the sensitivity of my main analysis to a particular forecasting method. Further, in addition to the GLS implied cost of equity, I compute and use the cost of equity based on the Fama–French (1993) 3-factor model (FFCOE) to check the robustness of my main results with the GLS implied cost of equity.[19]

2. Deregulation Indicator and Other Control Variables

To identify the effects of the 2003 amendment to the Worker Dispatching Act on the cost of equity on manufacturing firms relative to nonmanufacturing firms, I construct an indicator variable, DEREGULATION, which is equal to 1 for firms in the manufacturing sector in and after 2003, and 0 otherwise. The variable captures the treatment effect of the deregulation in 2003 and estimates a change in the cost of equity in manufacturing firms relative to nonmanufacturing firms.

Because I use manufacturing (nonmanufacturing) firms as the treatment (control) firms, I must ensure comparability between these two groups of firms. In terms of the differences in the fixed characteristics between them, I include firm fixed effects in all estimations to control for time-invariant differences between manufacturing and nonmanufacturing firms. Thus, I am left with controlling for the differences in the time-varying characteristics, which are relevant to the cost of equity, between the two groups of firms. I control for these differences by including a number of firm- and industry-level variables. In terms of the firm-level controls, I follow the guidance from the literature, such as Dhaliwal, Judd, Serfling, and Shaikh (2016), and include lagged values of i) IVOL, a firm's idiosyncratic stock volatility estimated from the Fama–French (1993) FFCOE using monthly returns in last 2 years; ii) ln(TA), the natural logarithm of total assets; iii) BE_ME, the book-to-market equity ratio, where market equity is equal to the number of outstanding shares multiplied by the stock price at a firm's fiscal year-end; iv) SGR, the growth rate of sales from the previous year; v) DEBT_TA, the total debt (i.e., short-term + long-term debt) over total assets; and vi) FA_TA, the net property, plant, and equipment over total assets. All accounting measures are winsorized at the 1st and 99th percentiles, if not stated otherwise, to avoid the effect of extreme values. Table A2 in the Supplementary Material presents the summary statistics of selected firm characteristics.

In addition to firm-level controls, I construct several industry-level variables to control for observable time-varying differences between the manufacturing and nonmanufacturing sectors. Specifically, I control for differences in industry-level

[18] See the Appendix for more details regarding the model-based forecasting method by Hou et al. (2012).

[19] All cost-of-equity estimates, including GLS and FFCOE, are winsorized at the 5th and 95th percentiles to avoid the effects of outliers.

sales, profitability, and the effects of past foreign exchange and stock market fluctuations. I include the lagged values of IND_LOG_SALES and IND_ROA, which are industry-year means of the natural log of firm-level sales and return on assets (ROA), respectively, to control for industry-level differences in fluctuations in sales and profitability. I compute the industry-year means at the 4-digit Japan Standard Industrial Classification (JSIC) level. The JSIC codes are analogous to the Standard Industrial Classification (SIC) codes in the United States. In my sample, there are 615 4-digit JSIC industries in total, out of which 350 4-digit JSIC industries belong to the manufacturing sector.

I also construct and include lagged values of M_JPY_USD, which is the annual change in the Japanese yen/US dollar exchange rate if a firm operates in a manufacturing industry, and 0 otherwise. Exchange-rate fluctuations could have a strong effect on manufacturing firms, which tend to export their products overseas. This variable controls for the differential effects of past exchange rates on the cost of equity between manufacturing and nonmanufacturing firms. In addition, I control for differences in past stock returns between manufacturing and nonmanufacturing firms, which could have driven the differences in the cost of equity between them, even in the absence of the deregulation. I include the lagged values of M_NM_RETURN, which is the mean of the annual stock returns among manufacturing (nonmanufacturing) firms in each year if a firm operates in a manufacturing (nonmanufacturing) industry.

IV. Empirical Analysis

A. Empirical Design and Identification Strategy: A DID Approach

In this section, I conduct a DID analysis to estimate the causal effects of the 2003 amendment to the Worker Dispatching Act on the cost of equity. My main specification takes the following form:

$$\begin{aligned}(1)\qquad \mathrm{COE}_{ij,t} \;=\;& \alpha+\beta \mathrm{DEREGULATION}_{j,t}+\gamma \mathrm{CONTROLS}_{i,j,t-1}\\ &+\mathrm{FIRM\_FE}+\mathrm{YEAR\_FE}+\varepsilon_{i,j,t},\end{aligned}$$

where i denotes a firm, j denotes a 4-digit JSIC industry, and t denotes a year. This empirical design is similar to those in recent articles on labor and capital structure (e.g., Matsa (2010), Simintzi et al. (2015), and Serfling (2016)). The dependent variable, COE, represents the cost of equity in percentage terms. In this section, it is GLS, my primary measure of the implied cost of equity by Gebhardt et al. (2001), computed with the cross-sectional forecasting model of Hou et al. (2012), as described in Section III.B.1. DEREGULATION is equal to 1 if a firm primarily operates in a 4-digit JSIC industry in the manufacturing sector in and after 2003, and 0 otherwise. The coefficient β measures the treatment effect of the deregulation in 2003 on the cost of equity for manufacturing firms relative to nonmanufacturing firms. Firm fixed effects absorb any other time-invariant differences between the treatment and control firms. Year fixed effects control for any common time-varying factors that affect the cost of equity of all firms.

Manufacturing firms and nonmanufacturing firms are, of course, different in observable and unobservable dimensions. I include a host of time-varying firm- and industry-level controls, as described in a previous section. In addition, I include different (linear) time trends for manufacturing and nonmanufacturing firms to control for a potential difference in linear trends in the cost of equity, even in the absence of the treatment. The clustered standard errors at the 4-digit JSIC industry level are calculated to account for the within-industry correlation of error terms because the deregulation applies to firms in specific industries (i.e., manufacturing industries). With this empirical design, I estimate within-firm changes in the cost of equity in manufacturing firms relative to nonmanufacturing firms in response to the deregulation in 2003.

B. The Effects of the Deregulation on the Cost of Equity: Main Results

I estimate equation (1) using the GLS implied cost of equity as the dependent variable. The variable of interest in equation (1) is the indicator variable DEREGULATION, which is equal to 1 for manufacturing firms in and after the 2003 deregulation year, and 0 otherwise. Table 3 presents the estimation results.

Column 1 of Table 3 includes DEREGULATION to estimate the effects of the 2003 amendment to the Worker Dispatching Act on the cost of equity in manufacturing firms relative to nonmanufacturing firms. As control variables, at a minimum, I include different time trends for manufacturing firms and nonmanufacturing firms to control for a potential difference in time trends in the cost of equity that are unrelated to the deregulation. In column 1, the coefficient of DEREGULATION is −0.297 and significantly negative, indicating that the cost of equity declined in manufacturing firms after the deregulation, relative to nonmanufacturing firms. In column 2, I add the 6 firm-level characteristics described in Section III.B.2 to the model in column 1: 1-year lags of the idiosyncratic stock volatility (IVOL), firm size (ln(TA)), book-to-market equity ratio (BE_ME), sales growth (SGR), leverage (DEBT_TA), and asset tangibility (FA_TA). The coefficient of DEREGULATION in column 2 remains significantly negative.

Because I compare changes in the cost of equity in response to the deregulation between manufacturing and nonmanufacturing firms, it is possible that some industry-level differences between those two groups of firms are responsible for my results. To alleviate this concern, in columns 3–5 of Table 3, I add several industry-level variables to control for some specific differences between manufacturing and nonmanufacturing firms that might affect the cost of equity. In column 3, I include 1-year lags of industry-year means of (the natural log of) sales (IND_LOG_SALES) and profitability (IND_ROA) at the 4-digit JSIC level. The coefficient of DEREGULATION is −0.269 and remains significant. Furthermore, in columns 4 and 5, I add two variables to control for differences in past stock returns and potentially different effects of past foreign exchange fluctuations on the cost of equity between manufacturing and nonmanufacturing firms. In column 4, I include a 1-year lag of M_NM_RETURN, which is the mean of annual stock returns among manufacturing (nonmanufacturing) firms in each year if a firm operates in a manufacturing (nonmanufacturing) industry. The coefficient of DEREGULATION is −0.338 and remains highly significant. In column 5, I include a 1-year lag of M_JPY_USD, which is the annual change in the Japanese yen/US dollar exchange rate if a firm operates in a manufacturing industry, and 0 otherwise. It turns out that accounting for the effects of past exchange

TABLE 3

The Effects of the Deregulation in 2003 on the Cost of Equity

Table 3 presents the regression results of the cost of equity on the deregulation in 2003. The sample consists of all listed firms except financials and utilities. The sample period is 2000–2006. The dependent variable is GLS, the implied the cost of equity of Gebhardt et al. (2001), estimated with the earnings forecasting model of Hou et al. (2012). DEREGULATION is an indicator variable that is equal to 1 for manufacturing firms in and after 2003, and 0 otherwise. Firm-level control variables include i) IVOL, a firm's idiosyncratic volatility estimated from the Fama–French (1993) 3-factor model using monthly returns in last 2 years; ii) ln(TA), the natural logarithm of total assets; iii) BE_ME, the book-to-market equity ratio; iv) SGR, the growth rate of sales from the previous year; v) DEBT_TA, the total debt (i.e., short-term + long-term debt) over total assets; and vi) FA_TA, the net property, plant, and equipment over total assets. Industry-level controls include i) IND_LOG_SALES, the industry-year means of the natural logarithm of sales; ii) IND_ROA, the industry-year means of net income over total assets; iii) M_NM_RETURN, the mean of annual stock returns of manufacturing (nonmanufacturing) firms in each year if a firm operates in a manufacturing (nonmanufacturing) industry; and iv) M_JPY_USD, the annual change in the JPY/USD exchange rate if a firm operates in a manufacturing industry, and 0 otherwise. Firm and year fixed effects (FE) are included in all models. I also include different time trends for manufacturing and nonmanufacturing firms. The model in column 6 uses the sample matched by the entropy-balancing method described in Section IV.B. Clustered standard errors at the 4-digit Japan Standard Industrial Classification (JSIC) industry level are calculated to account for within-industry correlations of error terms. The standard error of each coefficient is reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Dependent Variable: GLS (%)	Full Sample 1	Full Sample 2	Full Sample 3	Full Sample 4	Full Sample 5	Matched Sample 6
DEREGULATION$_t$	−0.297***	−0.275**	−0.269**	−0.338***	−1.085***	−0.968***
	(0.115)	(0.118)	(0.118)	(0.123)	(0.261)	(0.330)
IVOL$_{t-1}$		−4.696***	−4.686***	−4.713***	−4.700***	−4.547***
		(0.407)	(0.408)	(0.406)	(0.404)	(0.639)
ln(TA)$_{t-1}$		0.406**	0.355**	0.358**	0.367**	0.085
		(0.165)	(0.167)	(0.167)	(0.167)	(0.190)
BE_ME$_{t-1}$		0.403***	0.403***	0.403***	0.405***	0.363***
		(0.042)	(0.042)	(0.042)	(0.042)	(0.065)
SGR$_{t-1}$		−0.067	−0.088	−0.098	−0.112	0.181
		(0.152)	(0.153)	(0.154)	(0.154)	(0.158)
DEBT_TA$_{t-1}$		−1.077***	−1.032***	−1.043***	−1.044***	−1.596***
		(0.349)	(0.350)	(0.350)	(0.349)	(0.556)
FA_TA$_{t-1}$		−0.542	−0.526	−0.514	−0.508	−0.935
		(0.456)	(0.453)	(0.453)	(0.453)	(0.684)
IND_LOG_SALES$_{t-1}$			0.251**	0.248**	0.250**	0.045
			(0.113)	(0.112)	(0.112)	(0.199)
IND_ROA$_{t-1}$			0.046	0.132	0.200	−1.539
			(0.939)	(0.934)	(0.930)	(1.214)
M_NM_RETURN$_{t-1}$				0.933**	1.496***	1.532***
				(0.403)	(0.407)	(0.548)
M_JPY_USD$_{t-1}$					−1.778***	−1.526**
					(0.566)	(0.705)
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Different time trends	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	13,112	13,112	13,112	13,112	13,112	11,115

rate changes is important for the cost of equity of manufacturing firms. The coefficient of DEREGULATION in column 5 is −1.085, which is much larger in absolute terms than the results in previous columns. This result is not surprising because manufacturing firms produce tradable goods, and their values are exposed to exchange rate fluctuations, assuming that their foreign currency exposures are not perfectly hedged.

In terms of the economic significance of the deregulation on the cost of equity, the result in column 5 of Table 3 with a full set of control variables shows that my measure of the cost of equity declined by 1.085% in manufacturing firms relative to nonmanufacturing firms after the amendment to the Worker Dispatching Act in 2003. To gauge the economic magnitude of the deregulation in monetary terms for a typical manufacturing firm, I make a back-of-the-envelope calculation of the reduction in required compensation for shareholders per year by the deregulation. The median value of the market equity in manufacturing firms in 2002 was approximately 15.0 billion JPY, which was equivalent to 120 million USD, given the average daily JPY/USD exchange rate of 125 JPY/USD in 2002. Hypothetically, if the cost of equity was 1.085% lower in 2002, my approximate estimate suggests that the required compensation for shareholders per year would have been lower by 15.0 billion JPY (median market equity in manufacturing firms in 2002) $\times$ 1.085% = 162.75 million JPY (or 1.30 million USD) for a typical manufacturing firm. This figure corresponds to approximately 20% of the median net income of manufacturing firms with positive profits in 2002.

Finally, in column 6 of Table 3, I estimate equation (1) using a matched sample between the manufacturing and nonmanufacturing firms to ensure covariate balance between them. To be more precise, I employ entropy-balancing matching (Hainmueller and Xu (2013)). This method computes and assigns weights to the control-group data such that a specified set of covariate moments (e.g., mean, variance, skewness, etc.) in the control group closely matches that set in the treatment group.[20] Regarding covariates, I match on eight firm-level characteristics (six firm-level characteristics included in column 5 plus ROA and firm age) in the year prior to the deregulation (i.e., 2002).[21] I require that my entropy-balancing method find weights for the control-group data such that the sample means of the covariates in the control group (i.e., nonmanufacturing firms) match those in the treatment group (i.e., manufacturing firms) in 2002. Table A3 in the Supplementary Material presents the sample means of the treatment group (i.e., manufacturing firms) and the *matched* control group (i.e., nonmanufacturing firms) in the eight firm-level characteristics and shows that the matching in the sample means is almost exact.[22] Then, after completing matching in 2002, I use the same weights for all observations of the control firms in the whole sample period (2000–2006) and estimate equation (1) with the treatment-group data. The result in column 6 shows, similar to earlier results, that the cost of equity declined in the treatment group (i.e., manufacturing firms) relative to the *matched* control group (i.e., nonmanufacturing firms) after the deregulation.

Overall, the estimation results in Table 3 consistently show a significant decline in the cost of equity in manufacturing firms relative to nonmanufacturing firms after the deregulation in 2003. I note that the results are robust to an alternative method of estimating future earnings by the IBES consensus analyst forecasts when I compute the GLS implied cost of equity. In Section A4 of the Supplementary Material, I provide the details of my computation of the GLS implied cost of equity with the analyst forecasts and present estimation results using this alternative measure in Table A4 of the Supplementary Material. Further, in

[20]See Jacob, Michaely, and Müller (2019) for a recent application of entropy-balancing matching in corporate finance.

[21]Ideally, I would like to match a treatment firm to a control firm within the same industry to make the treated firms and control firms more comparable than with my approach. This approach is, unfortunately, not possible because the treatment and control firms are in different industries in my setting.

[22]Strictly speaking, for the control-group data, they are *weighted* sample means.

Section A5 of the Supplementary Material, I compute the cost of equity based on the Fama–French (1993) FFCOE and repeat the main analysis in Table A5 of the Supplementary Material. The estimation results with these alternative measures of the cost of equity show a similar decline in the cost of equity in manufacturing firms after the deregulation. These results are most consistent with the interpretation that the cost of equity decreased in manufacturing firms due to an expected increase in flexibility in labor costs and a decrease in the firms' exposure to systematic risk after the deregulation.

C. Addressing Econometric Concerns

In my DID framework, the key underlying assumption behind my identification strategy is that in the absence of the deregulation, the average changes in the cost of equity would have been the same for both the treatment and control groups. In this section, I conduct a number of analyses to gauge the plausibility of the parallel-trends assumption in my setting. First, I estimate and compare the pretreatment trends of the cost of equity between manufacturing and nonmanufacturing firms before the deregulation bill was passed in parliament in 2003. If there were no significant differences in trends during the pretreatment years, the parallel-trends assumption would likely be valid in my setting. Second, to rule out a concern that missing or unobservable differences in time-varying characteristics between manufacturing and nonmanufacturing firms might have driven a difference in the changes in the cost of equity between those firms, I exploit the cross-sectional variations in potential exposure to the deregulation *within* the manufacturing firms and repeat the main analysis. As I discuss later, this second approach enables me to fully control for the differences in the time-varying characteristics between manufacturing and nonmanufacturing firms.

1. Pretreatment Trends between Manufacturing and Nonmanufacturing Firms

Because I define the manufacturing (nonmanufacturing) firms as treatment (control) firms, which are not randomly assigned, a plausible concern is whether the parallel-trends assumption is likely to hold in my setting. It is possible that a pretreatment difference in the trends of the cost of equity between the two groups of firms, which existed even before the deregulation, might be responsible for my finding. For instance, if the cost of equity was declining in manufacturing firms relative to nonmanufacturing firms even before the deregulation, my main finding might simply reflect this pretreatment difference in trends that was unrelated to the deregulation.

To examine the validity of the parallel-trends assumption in my setting, I follow a standard procedure from the literature to check whether the pretreatment trends of the cost of equity between the two groups were different (e.g., Bertrand and Mullainathan (2003), Roberts and Whited (2013), and Serfling (2016)). Specifically, I examine the trends of the cost of equity of the two groups around 2003 and then check whether the two groups had similar (or different) annual changes in the cost of equity prior to 2003.

I replace DEREGULATION with four variables: $\text{DEREGULATION}^{2001}$, $\text{DEREGULATION}^{2002}$, $\text{DEREGULATION}^{2003}$, and $\text{DEREGULATION}^{2004}$. The first three variables are indicator variables and equal to 1 for manufacturing firms in 2001, 2002, and 2003, respectively, and 0 otherwise. DEREGULATION^{2004} is equal to 1 for manufacturing firms from 2004 onward, and 0 otherwise. The first three variables essentially enable me to estimate different coefficients of the year dummies from 2001 through 2003 between manufacturing and nonmanufacturing firms. The last variable captures the permanent effect of the deregulation in 2003 on the cost of equity of manufacturing firms after 2004. If there was a significant difference in the trends of the cost of equity between manufacturing and nonmanufacturing firms prior to the amendment to the Worker Dispatching Act in 2003, I would find significant coefficients of DEREGULATION^{2001} and DEREGULATION^{2002}, which would capture different year effects between the manufacturing and nonmanufacturing firms. To reliably estimate these coefficients in the pretreatment years before 2003, in this section, I expand my sample period from 2000–2006 to 1999–2007 (i.e., extend it by 1 year at the beginning and the end of my main sample period).

Table 4 presents the estimation results. The model in column 1 is analogous to the model in column 5 of Table 3, except that I exclude the different (linear) time trends for manufacturing and nonmanufacturing firms here because they will be highly collinear with the four indicator variables that I estimate for checking the potentially different (nonlinear) trends between manufacturing and nonmanufacturing firms. The result in column 1 shows that the coefficients of DEREGULATION^{2001} and DEREGULATION^{2002} are not significantly different from 0, indicating that the pretreatment trends of the cost of equity between

TABLE 4

Trends of the Cost of Equity of Manufacturing and Nonmanufacturing Firms around the Deregulation in 2003

Table 4 presents the regression results examining the trends of the cost of equity of manufacturing and nonmanufacturing firms around the deregulation in 2003. The sample consists of all listed firms except financials and utilities. The sample period is 1999–2007. The dependent variable is GLS, the implied cost of equity of Gebhardt et al. (2001), estimated with the earnings forecasting model of Hou et al. (2012). DEREGULATION^{2001}, DEREGULATION^{2002}, and DEREGULATION^{2003} are indicator variables that are equal to 1 for manufacturing firms in 2001, 2002, and 2003, respectively, and 0 otherwise. DEREGULATION^{2004} is an indicator variable that is equal to 1 for manufacturing firms in and after 2004, and 0 otherwise. Firm- and industry-level control variables are defined in Table 3. Firm and year fixed effects (FE) are included in all models. The model in column 2 uses the sample matched by the entropy-balancing method described in Section IV.B. Clustered standard errors at the 4-digit Japan Standard Industrial Classification (JSIC) industry level are calculated to account for within-industry correlations of error terms. The standard error of each coefficient is reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Dependent Variable: GLS (%)	Full Sample 1	Matched Sample 2
DEREGULATION^{2001}	−0.210	−0.289
	(0.204)	(0.269)
DEREGULATION^{2002}	−0.086	−0.026
	(0.156)	(0.235)
DEREGULATION^{2003}	−0.563***	−0.560***
	(0.162)	(0.168)
DEREGULATION^{2004}	−0.370**	−0.388**
	(0.163)	(0.177)
Firm controls	Yes	Yes
Industry controls	Yes	Yes
Firm FE	Yes	Yes
Year FE	Yes	Yes
No. of obs.	16,730	13,898

manufacturing and nonmanufacturing firms before the deregulation were not significantly different. The result suggests that the parallel-trends assumption is likely to be valid in my setting. In contrast, the coefficients of DEREGULATION^{2003} and DEREGULATION^{2004} are significantly negative, showing that the cost of equity in manufacturing firms decreased in 2003 and in 2004 and onward after the deregulation. Note that the coefficient of DEREGULATION^{2003} is -0.563 and that of DEREGULATION^{2004} is -0.370, such that the cumulative effect of the deregulation ($-0.563\% - 0.370\% = -0.933\%$) roughly matches the economic magnitude of DEREGULATION (-1.085%) in column 5 of Table 3.

In column 2 of Table 4, I estimate the same equation using the matched sample constructed in Section IV.B through an entropy-balancing methodology, and the result is very similar. Overall, the results in Table 4 are consistent with the interpretation that similar trends in the cost of equity in the treatment and control groups diverged in 2003 because the financial markets expected an increase in flexibility in labor costs after the deregulation, leading to lower operating leverage and cost of equity in manufacturing firms.

2. Cross-Sectional Variations within Manufacturing Firms

Although my analysis of pretreatment trends provides supportive evidence of the validity of the parallel-trends assumption in my setting, a concern might remain that some unobservable time-varying differences between manufacturing and nonmanufacturing firms are responsible for my results. First, it is possible that some unobservable demand or supply shocks specific to manufacturing or nonmanufacturing firms could have led to a difference in changes in the cost of equity between these two groups of firms. Second, some policy recommendations or regulation changes other than the one related to temporary agency workers might be responsible for my results. In fact, as I describe in Section II.A, the Council for Regulatory Reform, which was set up by the government in 2001, has issued several reports since its formation. Those reports contained other policy recommendations in addition to the recommendation on the deregulation of temporary agency workers. Thus, these recommendations, which were mostly related to nonmanufacturing sectors, could have affected the cost of equity in my control firms (firms in the nonmanufacturing sectors).[23] An ideal means of addressing the potential presence of these confounding factors would be to include a full set of different year fixed effects for manufacturing and nonmanufacturing firms, which is unfortunately not possible because they will also absorb the effects of the deregulation related to temporary agency workers on manufacturing firms.

To circumvent this problem, I exploit the cross-sectional variations in the firms' potential exposure to the deregulation *within* manufacturing firms. These variations should help me to identify the effects of the deregulation on a subset of manufacturing firms that had a high degree of exposure to the deregulation relative to the remaining manufacturing firms, which had a low degree of exposure to the deregulation, and nonmanufacturing firms. This empirical

[23]Some of those recommendations were related to a reform of ongoing business practices in nonmanufacturing sectors.

design is akin to a difference-in-difference-in-differences (DDD) estimation, in which the treatment group consists of manufacturing firms with a high degree of exposure to the deregulation, whereas the two control groups consist of manufacturing firms that had a low degree of exposure to the deregulation and nonmanufacturing firms. These variations are also useful because I can include a full set of different year effects for manufacturing and nonmanufacturing firms, such that I will be able to fully control for potential confounding events that might have affected manufacturing and nonmanufacturing firms differently.

To measure a firm's degree of potential exposure to the deregulation, first, I use the fact that the deregulation in 2003 applied only to temporary work agencies that dispatch their workers to manufacturing line work in firms located in Japan. This point implies that the deregulation should significantly affect manufacturing firms that produce goods and use labor primarily in production plants in Japan but not those that produce goods and use labor primarily in plants located in other countries. In other words, firms whose plants are diversified internationally would be less subject to the deregulation in Japan than would other firms producing goods mostly within Japan.

To estimate a firm's geographic distribution of manufacturing plants, I collect the segment information of manufacturing firms by geographic location. In reporting the results, the segments are divided into the domestic segment (i.e., Japanese segment) and foreign segments, where foreign segments are typically grouped by continent (e.g., North American segment, European segment, Asian segment, etc.). In the early 2000s in Japan, a firm was required to report the information for a foreign segment if the sales of that segment exceeded 10% of a firm's total sales. Approximately one-half of the manufacturing firms reported information for foreign segments that included some key accounting items, such as sales and total assets, by geographic location during my sample period. Those manufacturing firms that did not report information for foreign segments were supposed to have most of their operations located in Japan.

Ideally, I would like to obtain data on property, plant, and equipment by location, which would be a good proxy to capture the sizes of the manufacturing plants located in Japan and overseas for each firm. Unfortunately, the data on fixed assets are not available. However, the segment information reports the total assets by location. Assuming that the ratio of fixed assets to total assets in each location is roughly constant, I infer the relative size of a firm's domestic plants to all of its production plants by computing the ratio of a firm's total assets in the domestic segment to the sum of its assets in all geographic segments in 2002 (i.e., 1 year before the implementation of the deregulation in 2003). This ratio is my proxy for the relative size of a firm's domestic plants to all of its production plants; a higher value of the ratio would imply a higher degree of exposure to the deregulation. Then, I classify a firm into the group of high (low) exposure to the deregulation if the firm's ratio is higher (lower) than the median value of the ratios of all firms in the manufacturing sector. As mentioned, approximately one-half of manufacturing firms did not report information for foreign segments because those firms were likely to have most of their operations located in Japan. Thus, I also classify those firms that did not report information for foreign segments into the group of high exposure to the deregulation.

I conjecture that the deregulation would have stronger effects on firms in the group with high exposure than on those in the group with low exposure to the deregulation. Table 5 presents the estimation results. All models include firm and year fixed effects and different time trends for manufacturing and nonmanufacturing firms. The variable of interest is the interaction term between DEREGULATION and HIGH_EXPOSURE. The indicator variable HIGH_EXPOSURE is equal to 1 for manufacturing firms with high exposure to the deregulation, and 0 otherwise. Given that DEREGULATION is equal to 1 for all manufacturing firms in and after 2003, the interaction term captures the effects of the deregulation on the cost of equity of manufacturing firms with high exposure to the deregulation, relative to the other manufacturing firms and nonmanufacturing firms.

The model in column 1 of Table 5 includes DEREGULATION and DEREGULATION × HIGH_EXPOSURE, whereas firm- and industry-level controls are added in columns 2–4. Note that HIGH_EXPOSURE is a firm-level time-invariant variable and, thus, is absorbed by firm fixed effects. The coefficients of my variable of interest, DEREGULATION × HIGH_EXPOSURE, are significantly negative, showing that the cost of equity declined more in firms that had higher

TABLE 5

Cross-Sectional Effects of the Deregulation in 2003 on the Cost of Equity within Manufacturing Firms

Table 5 presents the regression results examining the cross-sectional effects of the deregulation in 2003 on the cost of equity in manufacturing firms, using the variations in potential exposure to the deregulation in 2003 within manufacturing firms. The sample consists of all listed firms except financials and utilities. The sample period is 2000–2006. The dependent variable is GLS, the implied cost of equity of Gebhardt et al. (2001), estimated with the earnings forecasting model of Hou et al. (2012). DEREGULATION is an indicator variable that is equal to 1 for manufacturing firms in and after 2003, and 0 otherwise. HIGH_EXPOSURE is an indicator variable that is equal to 1 for manufacturing firms with high exposure to the deregulation, and 0 otherwise. Firms with high exposure to the deregulation are defined in Section IV.C.2. HIGH_LABOR is an indicator variable that is equal to 1 for manufacturing firms with high labor intensity, and 0 otherwise. Firms with high labor intensity are defined in Section IV.C.2. Firm- and industry-level control variables are defined in Table 3. Firm fixed effects (FE) are included in all models. I also include different time trends for manufacturing and nonmanufacturing firms. Columns 1–5 include year FE, whereas column 6 includes different year FE for manufacturing and nonmanufacturing firms. Clustered standard errors at the 4-digit Japan Standard Industrial Classification (JSIC) industry level are calculated to account for within-industry correlations of error terms. The standard error of each coefficient is reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Dependent Variable: GLS (%)	1	2	3	4	5	6
$DEREGULATION_t$	0.064	0.030	0.041	−0.775***	−0.691**	
	(0.142)	(0.141)	(0.142)	(0.271)	(0.276)	
$DEREGULATION_t$ × HIGH_EXPOSURE	−0.451***	−0.382***	−0.387***	−0.388***	−0.384***	−0.384***
	(0.103)	(0.093)	(0.095)	(0.095)	(0.100)	(0.100)
$DEREGULATION_t$ × HIGH_LABOR					−0.144*	−0.144*
					(0.082)	(0.082)
$IND\_LOG\_SALES_{t-1}$			0.259**	0.258**	0.251**	0.252**
			(0.112)	(0.112)	(0.114)	(0.115)
IND_ROA_{t-1}			0.086	0.240	0.177	0.173
			(0.946)	(0.937)	(0.971)	(0.967)
$M\_NM\_RETURN_{t-1}$				1.498***	1.500***	
				(0.407)	(0.408)	
$M\_JPY\_USD_{t-1}$				−1.779***	−1.696***	
				(0.567)	(0.570)	
Firm controls	No	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	No
Different year FE	No	No	No	No	No	Yes
Different time trends	Yes	Yes	Yes	Yes	Yes	Yes
No. of obs.	13,112	13,112	13,112	13,112	12,883	12,883

exposure to the deregulation in the manufacturing sector. The result in column 4 suggests that the cost of equity declined by 0.388% more for firms with higher exposure to the deregulation.

Next, I further cement the validity of my causal interpretation by exploiting the variation in the labor intensity across firms within the manufacturing sector in columns 5 and 6 of Table 5. Because the deregulation pertains to temporary agency workers in production line work, the negative effects of the deregulation on the cost of equity should be particularly pronounced for firms with high labor intensity within the manufacturing sector. To test this conjecture, I construct the variable HIGH_LABOR to capture the variation in the labor intensity among manufacturing firms. I first compute a firm's labor–capital ratio, which is the number of employees over machinery and equipment, in 2002. Next, I classify a firm into the group of firms with high (low) labor intensity if the firm's labor–capital ratio is higher (lower) than the median value of the labor–capital ratios of all firms in the manufacturing sector. Then, HIGH_LABOR is equal to 1 for manufacturing firms with high labor intensity, and 0 otherwise.

I examine whether the negative effects of the deregulation on the cost of equity are more pronounced for firms with a higher labor intensity (i.e., HIGH_LABOR firms). Specifically, in column 5 of Table 5, I add the interaction term, DEREGULATION × HIGH_LABOR, to column 4.[24] The coefficient of this term captures the effects of the deregulation on the cost of equity of firms with high labor intensity in the manufacturing sector. The result in column 5 shows that DEREGULATION × HIGH_LABOR is −0.144 and significant at the 10% level, indicating that the cost of equity declined more in manufacturing firms with higher labor intensity after the deregulation.

Last, the cross-sectional variations in exposure to the deregulation within the manufacturing firms enable me to include a full set of different year fixed effects for manufacturing and nonmanufacturing firms in column 6 of Table 5, such that I can effectively control for any differences in time-varying confounding factors between manufacturing and nonmanufacturing firms. The different year fixed effects absorb DEREGULATION, M_NM_RETURN, and M_JPY_USD in column 6. The estimation result in column 6 shows that the coefficients of DEREGULATION × HIGH_EXPOSURE and DEREGULATION × HIGH_LABOR remain significantly negative, alleviating the concern that unobservable differences between manufacturing and nonmanufacturing firms are responsible for my main findings. Overall, the cross-sectional tests using the variations within manufacturing firms support my causal interpretation of the effects of the deregulation on the cost of equity.

D. Changes in the Number of Temporary Agency Workers and the Cost of Equity

My results in the previous sections consistently show that the cost of equity declined in manufacturing firms, particularly in those with high exposure and susceptibility to the deregulation, after the amendment to the Worker Dispatching Act was approved in the parliament in 2003. However, it is not yet clear whether the

[24]HIGH_LABOR is absorbed by firm fixed effects.

decline was truly driven by an (expected) increase in temporary agency workers in the manufacturing sector after the deregulation, although the overall number of temporary agency workers in manufacturing firms actually increased after the deregulation officially went into effect in 2004.

To better understand the actual mechanism behind a decrease in the cost of equity after the deregulation, I examine how the magnitude of the negative relationship between the deregulation and cost of equity differs with the size of a subsequent increase in temporary agency workers at the industry level *within* the manufacturing sector. Although the total number of temporary agency workers increased after the deregulation in the manufacturing sector, there was some variation in the size of the increases across manufacturing industries. For instance, the growth rates of temporary agency workers after the deregulation were higher in industries related to machinery production than were others within the manufacturing sector.

I use this cross-sectional variation in the manufacturing sector to examine the validity of my conjectured mechanism. If my findings in the previous sections truly reflect the effects of deregulation on the cost of equity through temporary agency workers, the negative relationship between the deregulation and cost of equity would be more pronounced in industries in which the number of temporary agency workers increased more after the deregulation. To test this conjecture, I collect data on the number of temporary agency workers in each 2-digit JSIC industry.[25] I calculate the growth rate of temporary agency workers from 2002 to 2004 for each 2-digit JSIC industry in the manufacturing sector. Next, I classify an industry into the group of industries with high (low) growth if the growth rate of temporary agency workers in the industry is above (below) the median value of the growth rates of all industries in the manufacturing sector. Then, I create an indicator variable, HIGH_GROWTH, which is 1 for firms in the group of high-growth industries, and 0 otherwise. Finally, I construct an interaction term between DEREGULATION and HIGH_GROWTH to examine whether the effects of the deregulation on the cost of equity were heterogeneous among the industries in the manufacturing sector, consistent with the cross-sectional variation among industries in subsequent increases in the number of temporary agency workers within the sector after the deregulation.

Table 6 presents the estimation results. If the cost of equity decreased due to an expected increase in temporary agency workers in the manufacturing sector after the deregulation, the magnitude of the decrease could be greater in industries in which the number of temporary agency workers increased more ex post in the sector. Columns 1–4 in Table 6 are analogous to those in Table 5, except that I replace HIGH_EXPOSURE with HIGH_GROWTH.[26] In columns 1–4, the coefficients of the interaction term are significantly negative, indicating that the cost of equity decreased more in industries in which the number of temporary agency workers increased more after the deregulation in the manufacturing sector, which

[25]The data are available from the Economic Census for Business Activity, conducted annually by the Ministry of Economy, Trade, and Industry. There are 24 2-digit JSIC industries in the manufacturing sector in my sample.

[26]HIGH_GROWTH is absorbed by firm fixed effects.

TABLE 6
Changes in the Number of Temporary Agency Workers and the Cost of Equity within Manufacturing Firms

Table 6 presents the regression results examining the heterogeneous effects of the deregulation in 2003 on the cost of equity in manufacturing firms, using variation in changes in the number of temporary agency workers between 2002 and 2004 within manufacturing industries. The sample consists of all listed firms except financials and utilities. The sample period is 2000–2006. The dependent variable is GLS, the implied cost of equity of Gebhardt et al. (2001), estimated with the earnings forecasting model of Hou et al. (2012). DEREGULATION is an indicator variable that is equal to 1 for manufacturing firms in and after 2003, and 0 otherwise. HIGH_GROWTH is an indicator variable that is equal to 1 if a firm belongs to an industry that had high growth in temporary agency workers between 2002 and 2004 in the manufacturing sector, and 0 otherwise. Industries with high growth in temporary agency workers are defined in Section IV.D. Firm- and industry-level control variables are defined in Table 3. Firm fixed effects (FE) are included in all models. I also include different time trends for manufacturing and nonmanufacturing firms. Columns 1–4 include year FE, whereas column 5 includes different year FE for manufacturing and nonmanufacturing firms. Clustered standard errors at the 4-digit Japan Standard Industrial Classification (JSIC) industry level are calculated to account for within-industry correlations of error terms. The standard error of each coefficient is reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Dependent Variable: GLS (%)	1	2	3	4	5
$DEREGULATION_t$	−0.171	−0.172	−0.160	−0.968***	
	(0.123)	(0.126)	(0.126)	(0.266)	
$DEREGULATION_t \times$ HIGH_GROWTH	−0.230**	−0.186**	−0.195**	−0.195**	−0.195**
	(0.099)	(0.089)	(0.091)	(0.091)	(0.091)
$IND\_LOG\_SALES_{t-1}$			0.251**	0.250**	0.250**
			(0.113)	(0.112)	(0.113)
IND_ROA_{t-1}			0.188	0.341	0.336
			(0.937)	(0.929)	(0.925)
$M\_NM\_RETURN_{t-1}$				1.502***	
				(0.410)	
$M\_JPY\_USD_{t-1}$				−1.756***	
				(0.564)	
Firm controls	No	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	No
Different year FE	No	No	No	No	Yes
Different time trends	Yes	Yes	Yes	Yes	Yes
No. of obs.	13,036	13,036	13,036	13,036	13,036

is consistent with my conjectured mechanism. In column 5, I include different year fixed effects for manufacturing and nonmanufacturing firms, and the result is essentially the same. The result in column 5 shows that the cost of equity declined by 0.195% more in industries in which the number of temporary agency workers increased more after the deregulation.

Overall, I find that the decline in the cost of equity was greater in firms in industries that had a greater increase in temporary agency workers in the manufacturing sector after the deregulation. These results are consistent with my conjecture that the cost of equity decreased due to an anticipated increase in temporary agency workers, which would increase the flexibility in labor costs and decrease the labor-induced operating leverage, after the deregulation.

E. The Effects of the Deregulation on Labor Leverage

Next, I conduct another analysis to check the validity of my conjectured mechanism for explaining the decline in the cost of equity in manufacturing firms. My conjecture is based on the premise that an increase in temporary agency workers in manufacturing firms would make the firms' labor costs more flexible after the deregulation than before, leading to a decrease in labor-induced operating leverage and the cost of equity. To cement the validity of my conjecture, I investigate whether the degree of labor leverage decreased in manufacturing firms, relative to nonmanufacturing firms, after the deregulation. Conceptually, the degree of labor leverage depends on the rigidity and the amount of the firms' labor expenses (Donangelo et al. (2019), Favilukis and Lin (2016a), (2016b), and Favilukis et al. (2020)). In this section, I specifically examine whether the flexibility (rigidity) in labor costs increased (decreased) in manufacturing firms after the deregulation. I also examine the effects of the deregulation on profit volatility because a change in the rigidity of labor expenses should be translated into a change in profit volatility, which could cause a change in the expected stock returns.[27]

I measure a change in the flexibility of labor costs in two ways. First, I use an approach that is similar in spirit to that of prior articles, such as those by Mandelker and Rhee (1984), Chen et al. (2011), Eisfeldt and Papanikolaou (2013), Serfling (2016), and Donangelo et al. (2019). I estimate the elasticity of labor costs to sales for all sample firms and examine whether the elasticity increased in manufacturing firms, relative to nonmanufacturing firms, after the deregulation. Further, if labor costs became more variable with sales, profits would vary less with sales, implying a decrease in the elasticity of profits to sales, or the degree of operating leverage, in manufacturing firms after the deregulation. A higher elasticity of labor costs to sales would be consistent with lower operating leverage and cost of equity.

I estimate the following equation using panel data of all public firms, except those in the financial and utilities industries, during the period of 1998–2008:

$$
\begin{aligned}
(2)\qquad \text{SALARIES\_GROWTH}_{i,j,t} &= \\
\alpha + \beta_1 &\text{SALES\_GROWTH}_{i,j,t} \\
&+ \beta_2(\text{AFTER\_2004}_t \times \text{SALES\_GROWTH}_{i,j,t}) \\
&+ \beta_3(\text{AFTER\_2004}_t \times \text{SALES\_GROWTH}_{i,j,t} \times \text{M\_DUMMY}_{j,t}) \\
&+ \beta_4(\text{AFTER\_2004}_t \times \text{M\_DUMMY}_{j,t}) \\
&+ \beta_5(\text{SALES\_GROWTH}_{i,j,t} \times \text{M\_DUMMY}_{j,t}) \\
&+ \text{FIRM\_FE} + \text{YEAR\_FE} + \varepsilon_{i,j,t},
\end{aligned}
$$

where i denotes a firm, j denotes a 4-digit JSIC industry, and t denotes a year. SALARIES_GROWTH is the rate of growth in labor expenses from the previous year. SALES_GROWTH is the rate of growth in sales from the previous year. AFTER_2004 is an indicator variable that is equal to 1 for observations in and after 2004 because the amendment to the act was approved in the parliament in June 2003 and went into effect in Mar. 2004. M_DUMMY is an indicator variable that is equal to 1 if a firm primarily operates in a 4-digit JSIC industry in the manufacturing sector. ε is an error term. I include firm and year fixed effects. Note that AFTER_2004 and M_DUMMY will be absorbed by these fixed effects.[28] I also include different time trends for manufacturing firms and nonmanufacturing firms to control for a potential difference in the time trends in labor-expense growth.

[27] In Section A6 of the Supplementary Material, I further investigate whether the deregulation had any effects on the amount of labor expenses relative to value added, that is, the labor share.

[28] I winsorize all growth variables at the 1st and 95th percentiles to avoid the effects of some abnormally high growth rates in labor costs, sales, or profits.

The variable of interest is the triple-interaction term between AFTER_2004, SALES_GROWTH, and M_DUMMY. Whereas the coefficient β_2 measures a change in the elasticity of labor costs to sales for all firms after the deregulation, the coefficient β_3 captures an incremental change in the elasticity in manufacturing firms, relative to nonmanufacturing firms, after the deregulation. If an increase in temporary agency workers after the deregulation in manufacturing firms made labor costs more flexible and sensitive to sales fluctuations, I would expect an increase in the elasticity in manufacturing firms after 2004, that is, a positive sign of β_3.

Table 7 presents the estimation results. The standard errors are clustered at the 4-digit JSIC industry level. In column 1, although the coefficient β_2 is not significant, the coefficient β_3 is significantly positive, suggesting that the elasticity of labor costs to sales increased in manufacturing firms, relative to nonmanufacturing firms, after the deregulation. The model in column 2 is similar to that in column 1, except that the dependent variable is PROFITS_GROWTH (i.e., the rate of growth in operating income from a previous year).[29] The elasticity of profits to sales measures the degree of operating leverage, and I examine whether the deregulation affected the degree of operating leverage in manufacturing firms. In column 2, the coefficient β_3 is significantly negative, indicating that the elasticity

TABLE 7

The Effects of the Deregulation on the Rigidity of Labor Expenses and the Degree of Operating Leverage

Table 7 presents the regression results examining the effects of the deregulation on the elasticity of labor expenses to sales in column 1 and on the elasticity of profits to sales (i.e., the degree of operating leverage) in column 2 in manufacturing firms, relative to nonmanufacturing firms. The sample consists of all listed firms except financials and utilities. The sample period is 1998–2008. The dependent variable in column 1 is SALARIES_GROWTH, the growth rate in labor expenses from the previous year. The dependent variable in column 2 is PROFITS_GROWTH, the growth rate in operating income from the previous year. I require the operating income in both the current year and the previous year to be nonnegative. SALES_GROWTH is the growth rate in sales from the previous year. AFTER_2004 is an indicator variable that is equal to 1 in and after 2004, and 0 otherwise, because the amendment to the act was approved in the parliament in June 2003 and went into effect in Mar. 2004. M_DUMMY is an indicator variable that is equal to 1 if a firm primarily operates in a 4-digit Japan Standard Industrial Classification (JSIC) industry in the manufacturing sector. Firm and year fixed effects (FE) are included in all models. I also include different time trends for manufacturing and nonmanufacturing firms. Clustered standard errors at the 4-digit JSIC industry level are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	SALARIES_ GROWTH	PROFITS_ GROWTH
Dependent Variable	1	2
SALES_GROWTH	0.351***	1.575***
	(0.031)	(0.131)
AFTER_2004 × SALES_GROWTH	−0.043	−0.020
	(0.030)	(0.146)
AFTER_2004 × SALES_GROWTH × M_DUMMY	0.141***	−0.733***
	(0.041)	(0.200)
AFTER_2004 × M_DUMMY	0.003	−0.045
	(0.006)	(0.038)
SALES_GROWTH × M_DUMMY	−0.090**	1.150***
	(0.039)	(0.167)
Firm FE	Yes	Yes
Year FE	Yes	Yes
Different time trends	Yes	Yes
No. of obs.	18,218	16,807

[29] To compute this variable, I require the operating income in the current year and the previous year to be nonnegative.

of profits to sales, or the degree of operating leverage, decreased in manufacturing firms, relative to nonmanufacturing firms, after the deregulation. These results are consistent with the interpretation that an increase in temporary agency workers decreased the rigidity of labor expenses and operating leverage in manufacturing firms after the deregulation.

Second, following Favilukis et al. (2020), I compute the standard deviations of labor-expense growth before and after the deregulation for each firm. If labor expenses became less rigid in manufacturing firms after the deregulation, I would expect an increase in the standard deviation of labor-expense growth after the deregulation in manufacturing firms. I would also expect a decrease in the profit volatility in manufacturing firms after the deregulation because a change in sales would have induced a greater change in labor expenses in those firms, making profits less volatile after the deregulation.

I estimate the following cross-sectional equation to examine whether a firm's standard deviation of labor-expense growth increased after the deregulation in manufacturing firms, relative to nonmanufacturing firms:

$$(3)\quad \Delta\text{INV\_SD\_SALARIES\_GROWTH}_{i,j} = \alpha + \beta\text{M\_DUMMY}_j + \varepsilon_{i,j},$$

where i denotes a firm, and j denotes a 4-digit JSIC industry. ΔINV_SD_SALARIES_GROWTH is the within-firm change in the *inverse* of the standard deviation of labor-expense growth from the pre-deregulation period (1998–2003) to the post-deregulation period (2004–2008).[30] A high value of the inverse term, or a low standard deviation of labor-expense growth, reflects high rigidity in the labor expense. Thus, an increase (decrease) in ΔINV_SD_SALARIES_GROWTH would imply an increase (decrease) in the labor-expense rigidity after the deregulation. M_DUMMY is an indicator variable that is equal to 1 if a firm primarily operates in a 4-digit JSIC industry in the manufacturing sector. ε is an error term. Heteroscedasticity-robust (Huber–White) estimates of the standard errors are reported in parentheses.[31]

Table 8 presents the estimation results. My variable of interest is M_DUMMY, which measures a within-firm change in the rigidity of labor expenses before and after the deregulation in manufacturing firms, relative to nonmanufacturing firms. In column 1, the coefficient of M_DUMMY is significantly negative, suggesting that labor costs became less rigid in manufacturing firms after the deregulation. In column 2, I estimate an analogous equation with a within-firm change in the standard deviation of profit growth before and after the deregulation, ΔSD_PROFITS_GROWTH, as the dependent variable.[32] If a reduction in the rigidity of labor expenses helped absorb the fluctuations in sales growth, the firm-level volatility in profit growth would have decreased after the deregulation in manufacturing firms. As expected, I find a negative coefficient of M_DUMMY, implying a decrease in profit volatility in those firms after the deregulation.

[30] I winsorize the inverse term in each period at the 1st and 99th percentiles to eliminate outliers.

[31] The subsequent estimation results with the standard errors clustered at the level of the 4-digit JSIC industry are qualitatively similar.

[32] I winsorize the standard deviation in each period at the 1st and 99th percentiles to eliminate outliers.

TABLE 8

Effects of the Deregulation on the Volatilities of Labor Expenses and Profits

Table 8 presents the regression results examining the effects of the deregulation on the volatility of labor expenses (column 1) and on the volatility of operating income (column 2) in manufacturing firms, relative to nonmanufacturing firms. The sample consists of all listed firms except financials and utilities. ΔINV_SD_SALARIES_GROWTH is a firm-level change in the *inverse* of the standard deviation of labor-expense growth from the pre-deregulation period (1998–2003) to the post-deregulation period (2004–2008). ΔSD_PROFITS_GROWTH is a firm-level change in the standard deviation of operating-income growth from the pre-deregulation period (1998–2003) to the postderegulation period (2004–2008). The dependent variables in columns 3 and 4 are the differences between those in columns 1 and 2 and ΔINV_SD_SALES_GROWTH and ΔSD_SALES_GROWTH, respectively, where ΔINV_SD_SALES_GROWTH is a firm-level change in the *inverse* of the standard deviation of sales growth from the pre-deregulation period (1998–2003) to the post-deregulation period (2004–2008), and ΔSD_SALES_GROWTH is a firm-level change in the standard deviation of sales growth from the pre-deregulation period (1998–2003) to the post-deregulation period (2004–2008). M_DUMMY is an indicator variable that is equal to 1 if a firm primarily operates in a 4-digit Japan Standard Industrial Classification (JSIC) industry in the manufacturing sector. Heteroscedasticity-robust (Huber–White) estimates of standard errors are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

	ΔINV_SD_ SALARIES_ GROWTH	ΔSD_ PROFITS_ GROWTH	ΔINV_SD_SALARIES_ GROWTH–ΔINV_SD_ SALES_GROWTH	ΔSD_PROFITS_ GROWTH–ΔSD_ SALES_GROWTH
Dependent Variable	1	2	3	4
M_DUMMY	−1.070*	−0.068***	−1.960**	−0.062***
	(0.628)	(0.016)	(0.893)	(0.016)
No. of obs.	1,638	1,506	1,638	1,506

In columns 3 and 4 of Table 8, the dependent variables are adjusted for the firm-level sales volatility that would directly affect the firm-level volatilities of labor expenses and profits. I construct ΔSD_SALES_GROWTH, a within-firm change in the standard deviation of sales growth from the pre-deregulation period (1998–2003) to the post-deregulation period (2004–2008). I also construct ΔINV_SD_SALES_GROWTH, a within-firm change in the *inverse* of the standard deviation of sales growth from the pre-deregulation period to the post-deregulation period. To adjust the dependent variables in columns 1 and 2 for the effects of sales volatility, I compute ΔINV_SD_SALARIES_GROWTH − ΔINV_SD_SALES_GROWTH and ΔSD_PROFITS_GROWTH − ΔSD_SALES_GROWTH and use them as the dependent variables in models in columns 3 and 4, respectively. The results in these models still show significant coefficients of M_DUMMY.

The results in Tables 7 and 8 show clear evidence that the flexibility in labor costs increased in manufacturing firms, relative to nonmanufacturing firms, after the deregulation, leading to lower operating leverage and profit volatility, consistent with my main findings of a decrease in the cost of equity in manufacturing firms after the deregulation. In Table A6 of the Supplementary Material, I examine the effects of the deregulation on the firm-level labor share, which is another determinant of labor leverage, and show that the labor share decreased in manufacturing firms, relative to nonmanufacturing firms, after the deregulation. With the findings taken together, I conclude that the degree of labor leverage decreased in manufacturing firms after the deregulation.

F. The Effects of the Deregulation on the Cost of Debt

Favilukis et al. (2020) argue that labor leverage should increase firms' credit risk because a higher rigidity of labor expenses makes firms more likely to default in bad times, especially when firms have a higher labor share. To support this argument, they show that labor share is positively associated with firms' credit risk, proxied by credit default swap (CDS) spreads and Moody's KMV Expected Default Frequency (EDF), and this positive relation is more pronounced when labor expenses are more rigid. If labor leverage affects firms' credit risk, it should also affect their cost of debt. In this section, I investigate the causal effects of labor leverage on the cost of debt, exploiting the labor-market deregulation as a quasi-natural experiment. Because the deregulation increased the number of temporary agency workers in manufacturing firms, which made their labor expenses more responsive to sales fluctuations after the deregulation, I posit that those firms would have experienced a decrease in the risk of default in bad times, leading to a lower cost of debt after the deregulation.

To examine the effects of the deregulation on the cost of debt, I employ an empirical design that is similar to the one used for the cost of equity. I estimate a panel-regression model in a DID framework and examine how the deregulation in 2003 affected the cost of debt in manufacturing firms, relative to nonmanufacturing firms. To measure a firm's cost of debt, I compute the yield to maturity of each corporate bond at the time of issuance and then take the difference between a bond's yield and a yield of a comparable Japanese Government Bond (JGB) of the same maturity (i.e., yield spread). Ideally, as in my analysis on the cost of equity, I would like to measure a change in the cost of debt by a price change within the same bond around the deregulation, which would enable me to control for any differences in fixed characteristics across different bonds. However, corporate bond markets are generally illiquid, and unfortunately, no reliable information on the actual trading prices of corporate bonds in over-the-counter (OTC) markets in Japan is available during my sample period. Thus, I use a corporate bond's yield spread at the time of issuance to proxy for a firm's cost of debt.[33]

To mitigate the concern that differences in bond yields reflect differences in firm-level characteristics, I construct and include several time-varying firm-level variables, as well as firm fixed effects, to control for time-varying and fixed heterogeneities across firms. I also include several variables to control for differences in issue-level characteristics, which I explain further in the following discussion. Thus, in this empirical design, I compare the yields of different bonds issued by the same firm before and after the deregulation and examine whether the yields of bonds issued after the deregulation were lower than those before the deregulation within the same firm.

I briefly describe my sample construction. I collect information on new corporate bond issues from a local financial data vendor, I-N Information Systems Ltd. I construct my sample from straight fixed-rate corporate bonds, with no option features, issued through public offerings or private placements between 1998 and 2008. I only consider corporate bonds issued by public firms that provide accounting and financial information. I exclude issuers in the financial and utilities industries. The resulting sample consists of 2,436 issues from 565 firms during the period of 1998–2008. For each bond issue, I collect several issue-level characteristics, such as the yield to maturity, the time to maturity, and the amount of

[33]I also explored ways to use the loan spreads of bank loans as a measure of the cost of debt. However, as far as I can determine, the data on loan spreads are missing for most of the loan deals made during my sample period in Japan in commercial databases, including DealScan.

proceeds, and construct several dummy variables indicating whether a bond is insured, whether a bond is backed by collateral, or whether a bond has covenants. I also construct several firm-level characteristics from the accounting and financial information, such as the stock return volatility, firm size, book-to-market equity ratio, ROA, sales growth, book leverage, and a firm's industry-level sales and ROA at the 4-digit JSIC industry level. The accounting and financial data are retrieved from the Nikkei NEEDS FinancialQUEST 2.0 database.

I estimate the following equation, which is similar to the one I estimate for the cost of equity, in a DID framework for the sample of newly issued corporate bonds during the period of 1998–2008:

$$
\begin{aligned}
(4) \qquad \mathrm{SPREAD}_{h,i,j,t} &= \alpha + \beta \mathrm{DEREGULATION}_{j,t} \\
&\quad + \gamma \mathrm{CONTROLS}_{h,i,j,t-1} + \mathrm{FIRM\_FE} \\
&\quad + \mathrm{YEAR\_FE} + \varepsilon_{h,i,j,t},
\end{aligned}
$$

where h denotes a bond issue, i denotes a firm, j denotes a 4-digit JSIC industry, and t denotes a year. The dependent variable, SPREAD, represents the difference in yields between a corporate bond and a JGB of the same maturity at the time of issuance in percentage terms. DEREGULATION is equal to 1 if a firm primarily operates in a 4-digit JSIC industry in the manufacturing sector in and after 2003, and 0 otherwise. The coefficient β measures the treatment effect of the deregulation in 2003 on the cost of debt on manufacturing firms relative to nonmanufacturing firms. I include the issue-level, firm-level, and industry-level control variables described previously. I also consider firm and year fixed effects. With firm fixed effects, I examine the within-firm changes in the yield spreads of corporate bonds issued by the same firm before and after the deregulation. Year fixed effects control for any common time-varying factors that affect the yield spreads of all bond issues. In addition, I include different time trends for manufacturing and nonmanufacturing firms to control for a potential difference in the linear trends in yield spreads, even in the absence of the deregulation. The clustered standard errors at the 4-digit JSIC industry level are calculated to account for within-industry correlation of error terms.[34]

Table 9 presents the estimation results. The model in column 1 includes firm- and industry-level control variables. The coefficient of DEREGULATION (−0.094) is significantly negative at the 10% level, suggesting that a firm's yield spread decreased by approximately 9 basis points (bps) in manufacturing firms, relative to nonmanufacturing firms, after the deregulation. In the model in column 2, I add several issue-level characteristics, and the coefficient remains significantly negative, indicating a decrease of 8.5 bps in the cost of debt. Overall, the results suggest a modest decrease in the cost of debt after the deregulation, compared with the decrease in the cost of equity reported in Table 3 (e.g., 108.5 bps in column 5). That the effect of the deregulation on the cost of debt is smaller than on

[34] I winsorize all accounting variables at the 1st and 99th percentiles and the yield spreads at the 1st and 95th percentiles to avoid the effects of some abnormally high spreads in my sample.

TABLE 9

Effects of the Deregulation on the Cost of Debt

Table 9 presents the regression results examining the effects of the deregulation in 2003 on the cost of debt. The sample consists of straight fixed-rate corporate bonds issued by public firms, excluding financials and utilities. The sample period is 1998–2008. The dependent variable is SPREAD (%), the difference in yields between a corporate bond and a Japanese Government Bond (JGB) of the same maturity at the time of issuance. DEREGULATION is an indicator variable that is equal to 1 for manufacturing firms in and after 2003, and 0 otherwise. Firm-level control variables include i) ln(TA), the natural logarithm of total assets; ii) BE_ME, the book-to-market equity ratio; iii) SGR, the growth rate of sales from the previous year; iv) DEBT_TA, the total debt (i.e., short-term + long-term debt) over total assets; v) ROA, the net income over total assets; and vi) SVOL, the stock return volatility, defined as the standard deviation of monthly stock returns in a calendar year. Industry-level controls include i) IND_LOG_SALES, the industry-year means of the natural logarithm of sales, and ii) IND_ROA, the industry-year means of net income over total assets. Issue-level characteristics include i) ln(MATURITY), the natural logarithm of time to maturity (in years); ii) ln(PROCEEDS), the natural logarithm of the amount of bond proceeds (JPY millions); iii) INSURANCE_DUMMY, a dummy variable that is equal to 1 if a bond is insured; iv) COLLATERAL_DUMMY, a dummy variable that is equal to 1 if a bond is backed by collateral; and v) COVENANTS_DUMMY, a dummy variable that is equal to 1 if a bond has covenants. Firm and year fixed effects (FE) are included in all models. I also include different time trends for manufacturing and nonmanufacturing firms. Clustered standard errors at the 4-digit Japan Standard Industrial Classification (JSIC) industry level are calculated to account for within-industry correlations of error terms. The standard error of each coefficient is reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Dependent Variable: SPREAD (%)	1	2
$DEREGULATION_t$	−0.094*	−0.085*
	(0.054)	(0.051)
$ln(TA)_{t-1}$	0.083	0.084
	(0.073)	(0.069)
BE_ME_{t-1}	0.004	0.005
	(0.003)	(0.003)
SGR_{t-1}	0.124	0.127
	(0.102)	(0.077)
$DEBT\_TA_{t-1}$	0.462**	0.613***
	(0.194)	(0.179)
ROA_{t-1}	−0.449	−0.285
	(0.592)	(0.588)
$SVOL_{t-1}$	0.406*	0.422*
	(0.214)	(0.219)
$IND\_LOG\_SALES_{t-1}$	−0.060	−0.043
	(0.070)	(0.064)
IND_ROA_{t-1}	−1.208*	−1.141*
	(0.652)	(0.612)
ln(MATURITY)		0.048**
		(0.021)
ln(PROCEEDS)		−0.028*
		(0.017)
INSURANCE_DUMMY		−0.303***
		(0.049)
COLLATERAL_DUMMY		−0.248***
		(0.075)
COVENANTS_DUMMY		0.052
		(0.032)
Firm FE	Yes	Yes
Year FE	Yes	Yes
Different time trends	Yes	Yes
No. of obs.	2,436	2,436

the cost of equity appears consistent with the differential impact of cash-flow risk on equity and debt. Higher labor leverage, or higher operating leverage in general, leads to higher upside and downside risks of cash flows, both of which would increase the equity risk premium. For debt, however, only the downside risk of cash flows in extremely bad times would matter to a firm's default risk and, thus, the cost of debt. Nevertheless, my results in Table 9 show a significant decrease in the cost of debt in manufacturing firms after the deregulation, consistent with the effects of labor leverage on credit risk documented by Favilukis et al. (2020).[35]

V. Conclusion

This article investigates whether firms' growing reliance on alternative work arrangements, especially temporary agency workers, affects their cost of equity capital. Because employers often cite high flexibility in labor adjustment through nonregular workers as a primary reason to increase their reliance on those workers, I posit that an increase in alternative work arrangements would reduce the operating leverage, which could affect the firms' cost of equity in two ways. On the one hand, a decrease in operating leverage would reduce firms' existing exposure to systematic risk, potentially leading to a decrease in the cost of equity (i.e., "direct effect"). On the other hand, a decrease in the labor-induced operating leverage could allow firms to take on more debt to take advantage of the greater benefits associated with debt financing, which could increase the equity beta and the cost of equity (i.e., "indirect effect"). Because these theories do not yield a unidirectional prediction, I empirically examine the total effect of alternative work arrangements on the cost of equity.

To estimate the causal effects of alternative work arrangements, I exploit a major labor-market deregulation in Japan as a quasi-natural experiment, which lifted a ban that had prohibited temporary agency workers from engaging in production line work in manufacturing. Using manufacturing (nonmanufacturing) firms as the treatment (control) group, I conduct a DID analysis and find that the cost of equity capital declined in manufacturing firms, relative to nonmanufacturing firms, after the parliament approved the deregulation in 2003. A number of robustness checks, including an analysis using variations only within manufacturing firms, suggest that the omitted or unobservable time-varying differences between manufacturing and nonmanufacturing firms are not responsible for my results. I also find that after the deregulation, the rigidity of labor expenses and the cost of debt decreased in manufacturing firms. Overall, the results are most consistent with the interpretation that the deregulation increased the flexibility in labor costs and decreased the labor-induced operating leverage, leading to lower exposure to systematic risk and a lower cost of capital in manufacturing firms after the deregulation.

Appendix. GLS Implied Cost of Equity Estimates by Hou et al. (2012)

I closely follow the forecasting model of firms' future earnings described by Hou et al. ((2012), pp. 506–508). For each year from 2000 through 2006, I estimate the following pooled cross-sectional regressions using the last 10 years of data:

$$\text{(A-1)} \quad \mathrm{E}_{i,t+\tau} = \alpha_0 + \alpha_1 \mathrm{TA}_{i,t} + \alpha_2 \mathrm{DIV}_{i,t} + \alpha_3 \mathrm{DD}_{i,t} + \alpha_4 \mathrm{E}_{i,t} + \alpha_5 \mathrm{NE}_{i,t} + \alpha_6 \mathrm{AC}_{i,t} + \varepsilon_{i,t+\tau},$$

[35]In Section A7 of the Supplementary Material, I also examine the effects of the deregulation on firm value. Table A7 in the Supplementary Material reports that the firm value increased in manufacturing firms, relative to nonmanufacturing firms, after the deregulation, consistent with a decrease in the overall cost of capital in manufacturing firms.

where $E_{i,t+\tau}$ denotes the earnings of firm i in year $t+\tau$ ($\tau = 1$–3); $TA_{i,t}$ is the total assets; $DIV_{i,t}$ is the dividend payment; $DD_{i,t}$ is a dummy variable that equals 1 for dividend payers, and 0 otherwise; $NE_{i,t}$ is a dummy variable that equals 1 for firms with negative earnings, and 0 otherwise; and $AC_{i,t}$ is accruals, proxied by net income minus operating cash flow. All variables except the dummy variables are in monetary terms (i.e., in JPY). I winsorize the earnings and other level variables each year at the 1st and 99th percentiles.

I use equation (A-1) to forecast future earnings. For each firm i and each year t in my sample, I compute earnings forecasts for up to 3 years into the future by multiplying the right-hand-side (RHS) variables as of year t with the coefficients from the pooled regression estimated using the last 10 years of data. For example, the computation of the GLS implied cost of equity in 2000 will typically require earnings forecasts for 2001 (1 year ahead), 2002 (2 years ahead), and 2003 (3 years ahead). To obtain the earnings forecast of firm i for 2001 (i.e., $\tau = 1$ in equation (A-1)), first, I estimate the following equation using all sample firms during the period of 1991–2000:

$$\text{(A-2)} \quad E_{i,t+1} = \alpha_0 + \alpha_1 TA_{i,t} + \alpha_2 DIV_{i,t} + \alpha_3 DD_{i,t} + \alpha_4 E_{i,t} + \alpha_5 NE_{i,t} + \alpha_6 AC_{i,t} + \varepsilon_{i,t+1}$$

for $t = 1991$–1999. Then, I use estimated coefficients ($\hat{\alpha}_0 \ldots \hat{\alpha}_6$) and the values of the RHS variables of firm i in 2000 to compute the firm's earnings forecast ($E^{\text{forecast}}_{i,2001}$) for 2001 as follows:

$$\text{(A-3)} \quad E^{\text{forecast}}_{i,2001} = \hat{\alpha}_0 + \hat{\alpha}_1 TA_{i,2000} + \hat{\alpha}_2 DIV_{i,2000} + \hat{\alpha}_3 DD_{i,2000} + \hat{\alpha}_4 E_{i,2000} + \hat{\alpha}_5 NE_{i,2000} + \hat{\alpha}_6 AC_{i,2000}.$$

A similar estimation procedure will apply to obtain the earnings forecasts of firm i for 2002 (i.e., $\tau = 2$ in equation (A-1)) and for 2003 (i.e., $\tau = 3$ in equation (A-1)).

At the end of June of each year t, I compute the GLS implied cost of equity for a given firm, which is essentially the internal rate of return that equates the end-of-June market equity to the present value of expected future cash flows. Following Hou et al. (2012), to ensure that the model-based earnings forecasts are based on information that is publicly available at the end of June of each year t, I impose a minimum reporting lag of 3 months. In other words, I match the end-of-June market equity in year t to the model-based earnings forecasts calculated for firms with fiscal year-ends (FYEs) between the end of April in year $t-1$ and the end of March in year t. For example, if a firm's FYE was at the end of March in 2000, I match the firm's end-of-June market equity in 2000 to the earnings forecasts calculated with the accounting information available at the end of March in 2000. I note that the end of March is the most common FYE in Japanese firms.

Supplementary Material

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