As California goes, so goes the nation? The impact of board gender quotas on firm performance and the director labor market*

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Abstract

On September 30, 2018, California became the first U.S. state to introduce a mandatory board gender quota for all firms headquartered in California. We find that the introduction of the quota is associated with significantly negative announcement returns for these firms. Consistent with the quota imposing frictions, the effect is larger for firms requiring more female directors to comply with the quota and for firms with poor corporate governance. We also document negative spillover effects to non-Californian firms. They are larger for firms operating in industries in which Californian firms lack more female directors, suggesting that valuable female directors may migrate from non-Californian to Californian firms. We also document negative spillover effects for firms headquartered in states that are more likely to follow California's lead. These are firms headquartered in states dominated by the Democratic Party and in states that legalized cannabis consumption. Finally, we show that, already as of month-end December 2018, female representation on the boards of Californian firms increased. Newly appointed female directors are younger, less experienced, and less independent than incumbent and leaving directors.

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1 Introduction

Women are still heavily underrepresented in leadership positions in the U.S. corporate sector. According to the Corporate Women Directors International (CWDI) 2018 report, women hold 21.4% of director positions on the boards of the Fortune Global 200 companies.[1] 19.5% of these companies have no female directors on their boards. In a country comparison, the U.S. places fifth in 2018 with 26.3% female directors. This number is surprisingly low, given that women in the U.S. account for about half of the employed population (Bertrand, Black, Jensen, and Lleras-Muney, 2018). Furthermore, the fraction of female directors in the U.S. is growing slowly, at an average rate of 0.5% annually. If no proactive measures are taken and the current rate of growth remains unchanged, it would take 48 years to achieve gender parity at U.S. boards.

In other countries with similar disparities, legislatures have responded to similar gender inequality by adopting mandatory board quotas. The first country to act was Norway, which introduced a gender quota of 40% female representation in 2003. Following Norway's lead, Belgium, France, Germany, Iceland, India, Israel, Italy, and Spain have all established similar quotas. These quotas vary in the fraction of women to be appointed, the set of firms that are subject to the quota, and the defined penalty for non-compliance.[2]

In the U.S., California is the first state to adopt a mandatory gender quota. On September 30, 2018, Governor Brown signed Senate Bill 826 into law. SB 826 requires that all national and foreign companies headquartered in California have at least one female director on their board by the end of 2019. Two female directors must be appointed to boards with five members, and three female directors must be appointed to boards with six members or more by the end of 2021. The statute is non-criminal, but penalties include a payment of $100,000 for the first violation, and $300,000 for each subsequent violation.

[1] See https://globewomen.org/CWDINet/.

[2] For an overview of international gender legislation, see Table 1.

It is still an open question whether the quota is illegal as it may conflict with the corporate internal affairs doctrine as well as California and U.S. federal civil rights laws.[3]

The effect of gender quotas on firm performance, board structure and even their effects on gender equality are still an open debate. While, in theory, board quotas can be an effective tool to improve gender equality, particularly if gender discrimination is the main obstacle for women to climb the corporate ladder (Bertrand, Black, Jensen, and Lleras-Muney, 2018), quotas have also been shown to lead to appointments of younger and less experienced women to corporate boards with adverse effects on firm performance (Ahern and Dittmar, 2012). The latter may result in a stigma put on so called "quota women", such that qualified women may still be reluctant to climb the corporate ladder even if quotas would help them to push through. Moreover, Bertrand, Black, Jensen, and Lleras-Muney (2018) find that seven years after the Norwegian quota fully came into effect, "it had very little discernible impact on women in business beyond its direct effect on the women who made it into boardrooms."

This paper uses the passage of the California senate bill 826 to investigate a number of questions related to the effect of gender quotas and passage of mandatory social laws applicable to corporations. The first question we explore is how the introduction of a mandatory gender quota affects Californian firms' valuations. We compute abnormal stock returns for firms headquartered in California and a matched group of control firms for different event windows surrounding the days of the gender quota's adoption and announcement in California. We observe a robust and significantly negative valuation effect on stock returns of firms affected by the quota. Specifically, firms headquartered in California have a 0.45% lower announcement return on the first day after the quota announcement than a group

[3] The internal affairs doctrine mandates that matters involving corporate boards are subject to the laws of the state of incorporation, not where the headquarters are located. By targeting all firms headquartered in California, the law deviates from the internal affairs doctrine. Moreover, Senate Bill 826 prioritizes one gender over the other as well as gender over other aspects of diversity. Hence, the law may violate, among others, civil rights laws.

of control firms headquartered in other U.S. states or the D.C. matched on size and industry. These results translate into a value loss of around 60 million USD on average per California-headquartered firm relative to non-California-headquartered firms. This value loss increases in economic magnitude for larger event windows and is robust to a number of alternative specifications.

The documented negative announcement effects are stronger for California-headquartered firms requiring more female directors to comply with the quota and for firms with lower corporate governance standards. This is consistent with the theory that smaller firms and firms with inferior corporate governance could face more significant obstacles to recruiting qualified female candidates. In contrast, we find the negative announcement effect to be muted for Californian firms with access to a larger network of female directors, which may facilitate the appointment of female directors and thus compliance with the quota. We also find that the market reaction to the quota is more positive (or less negative) the higher the firm's sustainability score. This positive relationship between sustainability and abnormal returns to the quota's announcement applies to both California-headquartered and non-California-headquartered firms, but is stronger for the latter. One conclusion that could be drawn from these results is that part of the market reaction around the quota announcement is caused by investors' disapproval of California's willingness to legislate non-economic values on California firms.

We further examine whether there are negative spillover effects to firms that are headquartered in states that are likely to follow California's lead. In recent years, California has frequently been the first state to enact progressive legislation that was later adopted by other states in the U.S. as well. For example, California pioneered legislation to promote reductions of carbon emissions from road vehicles introducing a federal waiver to set its own emissions standards, along with a zero-emission vehicle mandate. Thirteen states subsequently adopted California's stricter emissions standards.[4] Another example of California's leadership in legislative issues is the legalization of cannabis usage. It is reasonable to assume that California's actions on a mandatory gender quota may increase the likelihood that certain other states will follow.

We adopt multiple strategies to examine spillover effects on firms that are not headquartered in California. First, we examine firms headquartered in democratic states, which are arguably more likely to follow Californian legislation than republican states. In fact, we find spillover effects to be significantly larger in democratic states, and significantly reduced in republican states. This finding can be interpreted as investors assessing a higher likelihood of mandatory gender quotas being enacted in democratic states. Similarly, we find that firms headquartered in states that legalized cannabis react more negatively to the quota's introduction, a result that is also consistent with the perception of these firms being more likely to face such a quota themselves. Finally, we document negative spillover effects for firms that operate in industries in which Californian firms need to appoint a large number of female directors to comply with the quota. This finding supports the view that the gender quota is value-reducing because it increases the competition for scarce, experienced female directors.

Finally, we examine whether the board composition of Californian firms has changed in response to the new gender quota. We observe that, relative to control firms California-headquartered firms significantly increased female board representation by 0.45 percentage points three months after the introduction of the quota law (i.e., as of month-end December 2018). Californian firms that are under more pressure to fulfill the quota, i.e., those that require one (two) female director(s) to comply with the quota, have reacted more quickly to appoint female directors than a sample of control firms headquartered in other states.

[4] Known as the "Section 177" states, those 13 states are: Connecticut, Delaware, Maine, Maryland, Massachusetts, New Jersey, New Mexico, New York, Oregon, Pennsylvania, Rhode Island, Vermont, and Washington. See https://mde.maryland.gov/programs/air/mobilesources/pages/states.aspx.

On average, these firms increased female board representation by 0.58 (0.69) percentage points three months after the introduction of the quota. Some preliminary evidence on the skill set and experience of the newly appointed female directors indicates that they are younger, significantly less likely to possess industry experience or experience as (an outside) director of another listed firm and less likely to be independent than incumbent female and male directors.

Previous evidence on the valuation impact of mandatory gender quotas is ambiguous. While Ahern and Dittmar (2012) document significantly negative announcement returns for firms affected by the introduction of a gender quota in Norway, Eckbo, Nygaard and Thorburn (2018) suggest a value-neutral effect of the Norwegian gender quota on affected firms. The reason for these differing results may be that it is very difficult to establish a proper control group to which affected firms in Norway could be compared (Ferreira, 2015). As all public limited liability firms (i.e., ASA firms) in Norway were affected by the gender quota, control firms would have to be defined as either private firms (that are structurally different from ASA firms), or firms from neighboring countries or the U.S. We circumvent this problem by examining the impact of a gender quota within one country, i.e., the U.S. where only firms in one state are affected by the quota, while all other firms headquartered in different states can potentially serve as control observations.[5]

Our results are in line with Ahern and Dittmar (2012) in that they show a significantly negative short-term valuation effect of the mandatory gender quota. This finding is subject to varying interpretations. One is directly related to the gender quota itself. The value reductions we find may be attributable to investor assessment that the law will lead to the appointment of less-qualified directors and subsequent firm underperformance. In our

[5] In a contemporaneous paper, Hwang, Shivdasani, and Simintzi (2019) confirm the negative valuation impact of the Californian gender quota on firms affected by the quota even if no matching approach is used. They conjecture, but due to data limitations cannot directly test, that inelastic supply of qualified women contributes to the negative valuation effect. Our data allow for a more detailed analysis of changes in the director labor market in response to the gender quota.

tests, we find evidence that this may indeed be the case: We document negative spillover effects for firms in industries where competition for female directors is likely to be more intense. We also find that firms appoint younger and less experienced female directors to corporate boards in response to the mandatory quota.[6]

On the other hand, the strong negative valuation results we find are remarkable in that the law only requires the appointment of additional directors, not the removal of male directors. Thus, it is surprising that the simple addition of one to three directors would change firm value so substantially. A second interpretation of the results more in line with this skepticism is that the investor reaction is related to an assessment of the willingness of California (and other similarly politically aligned states) to impose non-economic legislation on firms headquartered in that state. Our results on the effect of this law on smaller firms, those with low corporate governance standards, and those with low sustainability scores imply that firms which might be most affected by future legislation react more strongly, which gives credence to this hypothesis. However, our findings on spillover effects highlight that the gender quota itself has had some effect and that perhaps both effects are at work here.

Under either hypothesis, our findings suggest that non-economic laws with respect to the corporation have economic effects. In the case of gender equality, studies have found that proactive measures to increase female representation on all management levels within a firm seem to be crucial to achieve a sustainable increase of the fraction of women in leadership.[7] Gender quotas at the board level do not find this evidential support and

[6] This finding is consistent with evidence in Ahern and Dittmar (2012) on the Norwegian quota, but contradicts the results of Bertrand, Black, Jensen, and Lleras-Muney (2018) and Ferrreira, Ginglinger, Laguna, and Skalli (2018), finding that quotas lead to the appointment of more qualified female directors in Norway and France, respectively.

[7] For example, slow career progression in the five years after the first childbirth has been shown to be a major obstacle for women in business (Keloharju, Knüpfer, and Tag, 2018). A proactive measure that mitigates this problem is government provided child care which allows women to preserve the value of their human capital in child bearing years and helps them pursue their careers in spite of starting a family (Chaochharia, Ghosh, Niessen-Ruenzi, and Schneider, 2019).

appear to be beneficial only to those women that are directly promoted to a board seat due to the quota, but not to all other women in business who were not appointed to boards (Bertrand, Black, Jensen, and Lleras-Muney, 2018).

2 The gender quota in California: Senate Bill 826

The Norway gender quota was introduced in 2003 and was followed by similar legislation in other countries, primarily in Europe. For example, Germany, France, and Italy established minimum gender quotas of up to 40 percent (see Table 1). In many of these countries, the adoption of mandatory quotas was preceded by precatory resolutions which called for companies to voluntarily increase female board representation.

In September 2013, California Senate Concurrent Resolution 62 was passed by both houses of the California state legislature. This resolution was non-binding and called for "every publicly held corporation in California with 9 or more director seats have a minimum of 3 women on its board, every publicly held corporation in California with 5 to 8 director seats have a minimum of 2 women on its board, and every publicly held corporation in California with fewer than 5 director seats have a minimum of one woman on its board." California was the first state in the U.S. to adopt such a non-binding resolution. Legislatures in Illinois (May 2015), Massachusetts (October 2015), Ohio (April 2016), Colorado (March 2017), and Pennsylvania (April 2017) all subsequently passed similar resolutions calling for an increase in the fraction of female directors on corporate boards in their states.

The three-year time frame specified in the California resolution ended on December 31, 2016. According to our own estimations, approximately 194 firms (44%) of the firms included in the Russell 3000 Index and headquartered in California failed to comply with the resolution's targeted number of female directors on that date. Among the 446 publicly traded California-headquartered firms included in the Russell 3000 index, female directors held 526 (14%) of seats, men held 3,090 (86%) of seats, and 126 (28%) firms had no female directors.

Senate Bill 826 was introduced on January 3, 2018 with a stated purpose of addressing the continued deficit in female directors at California publicly traded firms. The bill introduced graduated requirements for female directors applicable to any publicly held domestic or foreign corporation with its principal executive offices, according to the corporation's SEC 10-K form, located in California.[8] The Bill generally requires between 25% (1 / 4) and 50% (3 / 6) of female directors on the board. More specifically, the Bill requires that these California firms have a minimum of one female director on its board by the end of 2019. The Bill also requires that, by the end of calendar year 2021, California companies with five directors be required to have two female directors, and companies with six or more directors are required to have three female directors.

The Bill provides an enforcement mechanism which applies financial penalties to companies which do not comply. For a first violation, a fine of $100,000 is imposed; for a second or subsequent violation, the fine is increased to $300,000. A violation is defined as a "director seat required by this section to be held by a female, which is not held by a female during at least a portion of a calendar year." The maximum fine imposable under the requirements of the Bill is $900,000 per year, the fine imposable for an all-male board with six or more directors. While this may be economically insignificant for most public firms, we believe that there will be substantial pressure from institutional shareholders and other constituencies for boards to comply with the law rather than simply pay the penalty. This pressure would be applied more towards larger companies and those with high institutional ownership. For example, when Twitter had its initial public offering it had no

[8] The Bill defines a female director as "an individual who self-identifies her gender as a woman, without regard to the individual's designated sex at birth."

female directors and faced significant criticism; Twitter now has three female directors.[9]

Senate Bill 826 was introduced into the Senate on January 3, 2018 and had its first hearing on May 7, 2018. The Bill first passed the Senate on May 31 2018 (23:8:9 votes) and the Assembly on August 29, 2018 (41:13:26 votes), and the Senate passed the amendments made by the Assembly on August 30, 2018 (23:9:8 votes).

The Bill was presented to the Governor's office for signing on September 10, 2018. Governor Brown is known for his "willingness" to veto bills and vetoed 12% of bills in 2017 and 15% in 2016.[10] He did not initially announce what he would do with Senate Bill 826 and news and commentary at the time highlighted that it was uncertain whether he would sign the Bill.[11] On Sunday September 30, 2018, SB 826 was signed into law by Governor Brown and announced the same day.

Governor Brown signed the Bill despite significant commentary that it could be unconstitutional (Grundfest, 2018). California's own legislative analysis concludes that "the use of a quota-like system, as proposed by this bill ... may be difficult to defend." The argument for unconstitutionality has thus far focused on two grounds. First, that the Bill violates the internal affairs doctrine since it purported to regulate firms incorporated in

[9] See Claire Cain Miller, Curtain Is Rising on a Tech Premiere With (as Usual) a Mostly Male Cast, The N.Y. Times, Oct 4, 2013.

[10] See Angela Hart, Jerry Brown consistently signs more bills than GOP governors, Sacramento Bee, Oct. 17, 2017. See also Tim Arango and Jose A. Del Real, 5 Takeaways from California Gov. Jerry Brown's Last Bill Signing Session, The N.Y. Times, Oct. 1, 2018 (Governor Brown has a "willingness to wield the veto pen").

[11] See David Matsa, Norway's quota for women on boards suggests an overlooked benefit — to workers, Quartz, Sept. 28, 2018, available at https://qz.com/work/1406503/if-california-mandates-a-gender-quota-for-company-boards-it-could-be-good-for-workers/ ("Governor Brown has not indicated whether he will sign the bill."); Jorge L. Ortiz, Gender quotas: California ponders breakthrough bill to boost female executives, USA Today, Sept. 18, 2018 ("As California Gov. Jerry Brown ponders whether to sign a landmark bill . . . .").

Delaware or another State and headquartered in California.[12] Second, that the imposed quota violates the U.S. and California Constitutional equal rights protections. In his signing statement, Governor Brown stated “[there] have been numerous objections to this bill and serious legal concerns have been raised.” Nonetheless, Governor Brown asserted that “it’s high time corporate boards include the people who constitute more than half the ‘persons’ in America.”

3 Data and sample selection

3.1 Sample selection and data sources

To compile our sample, we first select all firms in Compustat with a data entry within one calendar year before September-end 2018, the date of the introduction of the gender quota in the state of California. We drop utility and financial firms (SIC codes 4940-4949 and 6000-6999, respectively), firms with missing information on the state in which they are headquartered, firms headquartered outside the US, firms with negative book value of equity, and firms with missing financial control variables as described below. Moreover, we drop firms that only list American Depository Receipts (issue IDs 90 or above) and firms without a listing on NYSE, AMEX, or NASDAQ (stock exchange codes 11, 12, and 14, respectively). This initial sample selection results in a sample of 2,562 firms that enter the sample with the most recent fiscal year-end balance sheet data in Compustat that predates September 30, 2018. 475 firms are headquartered in the state of California and the other 2,087 firms in one of the other 49 U.S. states or the D.C.

[12] In anticipation of the possibility that the bill will be revised to be in accordance with the internal affairs doctrine, we would expect the subset of firms that are both headquartered and incorporated in California to show the most negative announcement returns. However, while 455 firms in our sample are headquartered in California, only 17 firms are incorporated in California. 15 of the 17 firms incorporated in California are also headquartered in California and only two are not. Hence, the small number of observations and high correlation between California incorporation and headquarter location do not lend themselves to empirical tests.

To analyze the market reaction to the introduction of the gender quota, we supplement our sample with stock return data from Compustat.[13] Compustat reports more than one return series for 172 firms (6.7% of the sample). In these cases, we chose the time-series with the highest market capitalization as of the event date among those return series with sufficient data to run the market model.

Governor Brown signed the law on Sunday, September 30, 2018. On the same day, the adoption of the law was publicly announced. Hence, the first trading day, and thus the event date, is Monday, October 1. We compute daily abnormal returns (ARs) for a symmetric five-day window around the event date. Abnormal returns are calculated as the observed return less the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21, i.e., six trading days before the event date and four trading days before the start of the event window. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. For a firm to be included in our final sample, we require at least 125 daily return observations during the estimation window and complete return data during the entire five-day event window. We compute a set of alternative cumulative abnormal returns (CARs) over different sub-periods within the five-day event window. Our base case measure of the market's reaction to the introduction of the gender quota is a two-day event window, which includes the event day (October 1) and the following day (October 2). All abnormal return measures are winsorized at the 1st and 99th percentiles to mitigate the effect of outliers. We are unable to compute abnormal shareholder value

[13] At the time of writing, daily stock return data from the Center for Research in Security Prices (CRSP) are only available through the end of June 2018 and the update schedule provided by the Wharton Research Data Service (WRDS) suggests that daily stock return data from CRSP covering our time-period will not become available until August 2019. Moreover, the CRSP-Compustat matching table necessary to merge CRSP and Compustat is only available through January 2018. In order to analyze whether there are any material differences between CRSP's and Compustat's return data, we compare return observations in January 2018, the one month in our estimation window for which matched data are available from both databases. We find that only 298 out of 43,349 return observations (0.7%) differ by more than 0.01 percentage points, suggesting that return data in CRSP and Compustat are very similar.

changes around the announcement of the quota's adoption for 100 firms.

Most of our tests require detailed data on the board of directors of our sample firms. We obtain board data from BoardEx. BoardEx includes detailed information on current and past employments, education, involvement in non-profit organizations and club memberships, among other things, on almost all directors serving on boards of publicly listed U.S. firms. BoardEx is updated daily and organizes its data at the director level. This allows us to observe a firm's board structure at any point in time and, thus, to analyze how board structure changes in response to the introduction of a gender quota. Moreover, we can track individual directors over time and across firms. This enables us to construct detailed measures of directors' work experience and to construct measures of directors' connections to other directors or top executives (e.g., Custódio, Ferreira, and Matos, 2013; Engelberg, Gao, and Parsons, 2013). Applying all these filters results in a final sample of 2,455 firms, out of which 455 are headquartered in California and 2,000 are headquartered in other states of the U.S. or the D.C.

3.2 Descriptive statistics

Panel A of Table 2 reports descriptive statistics for the sample of 2,455 firms. 455 firms (18.5%) have their headquarters in California. On average, sample firms have total assets of almost USD 6.4 billion, carry 22.1% of assets as long-term debt or debt in current liabilities, have 22.1% of assets invested in property, plant, and equipment, invest 8.4% of total assets per year in research and development, and achieve a return on assets of -1.9%.[14]

[14] The return on assets (ROA), defined as operating income before depreciation and amortization scaled by total assets, is surprisingly low when compared to the median of 8.6%. As ROA is winsorized at the 1st and 99th percentiles, this suggests that the sample is comprised of a substantial number of firms with strongly negative ROA values. In unreported tests, we find this pattern not to be an artifact of our sample selection or variable construction. Specifically, we find similar numbers when computing ROA as net income scaled by total assets (mean: -9.2%; median: 1.9%), when using EBITDA scaled by total assets (mean: -1.9%; median: 8.6%), and when winsorizing ROA at the 5th and 95th percentiles (mean: -0.1%; median: 8.6%). Moreover, we find even more negative mean ROA values when looking at the entire (winsorized) cross-section of Compustat firms (mean: -31.4%; median: 4.1%).

As of September 30, our sample firms have a mean (median) board size of 8.1 (8.0), out of which 15.3% (14.3%) are female directors. The quota imposed on firms headquartered in California mandates that firms have to have, by the end of the calendar year 2021, three female directors if board size is six or more, two female directors if board size is five, and one female director if board size is four or less. Using data on board size and directors' gender, we find that, at the time of the adoption of the quota law, 84.0% of all firms are not in compliance with the mandated quota that would apply to the current board size (2021 requ. failed (d)).[15] To comply with the quota, firms would have to appoint on average 1.5 (median: 2.0) female directors (# female directors missing), or, if expressed as a fraction of board size, would have to increase female board representation on average by 21.4% (median: 22.2%) to comply with the gender quota (Shortfall (%)).

3.3 Construction of matched control sample

In most our tests, we compare the reaction to the quota of treated firms, i.e., firms headquartered in California, with the reaction of a sample of control firms, i.e., firms headquartered in the other 49 U.S. states or the D.C. The validity of such tests rests on the assumption that the two groups of firms are comparable before the onset of treatment. Prior research looking at the Norwegian gender quota, which affected all public limited companies (ASA) in Norway, had to resort to private limited liability companies (e.g., Matsa and Miller, 2013; Eckbo, Nygaard, and Thorburn, 2018), listed firms from other Scandinavian countries (e.g., Ahern and Dittmar, 2012; Matsa and Miller, 2013), or listed U.S. firms (e.g., Ahern and Dittmar, 2012) as a control group. Both, foreign and privately held firms, constitute imperfect control groups as they are subject to, for example, different corporate governance (guidelines) and regulation. On top of that, foreign firms even differ in terms of the macroeconomic environment (Ferreira, 2015). A major advantage of our

[15] Note that we compute this and the following variables related to the quota compliance for all sample firms, even though only firms headquartered in California are directly affected by the quota.

setting is that we can draw control firms from a large pool of listed domestic companies with headquarters in another U.S. state or the D.C.

To obtain treatment and control groups that are similar in terms of observable financial and in particular board characteristics, we apply alternative matching procedures to define a sample of control firms. Our first matching procedure, which serves as a base case, consists of choosing, for each of our 455 treatment firms headquartered in California, three non-California-headquartered firms that share the same primary two-digit SIC and are closest in terms of total assets. We match with replacement, i.e., a firm in the control sample may serve as a matched control firm to more than one treatment firm, but we include every control firm only once in the sample. The resulting final sample comprises 1,232 firms, 455 in the treatment group and 777 in the control group.[16]

Balancing tests reported in Table OA.1 in the Online Appendix show that this matching procedure indeed results in a control sample that is similar to the treatment sample in terms of board characteristics. Specifically, the results in Panel A, which compares California-headquartered to all non-California-headquartered firms in the full sample, show that California-headquartered firms are not only significantly different from non-California-headquartered firms in terms of financial control variables, but also have smaller boards and require more female directors to comply with the quota requirements. Panel B of Table OA.1 compares California-headquartered firms to the matched control sample. While differences in financial controls are reduced, all but firm size, which is a matching characteristic, remain significant at conventional levels. More importantly, however, differences in board structure and requirements to comply with the quota turn economically and statistically insignificant. To address concerns of differences in financial controls between

[16] To provide an example, our matching procedure assigns the three biopharmaceutical firms Abbvie Inc. (headquarter: IL), Merck & Co. (NJ), and Abbott Laboratories (IL) as control firms to Amgen Inc., a biopharmaceutical firm headquartered in Thousand Oaks, CA. All four firms are active in the same two-digit SIC code industry 28 (Chemical & Allied Products). Differences in total assets of the three control firms relative to Amgen Inc. amount to -11.5%, 9.9%, and -4.6%, respectively.

California-headquartered firms and non-California-headquartered firms, we conduct robustness tests in which we control for these observable financial firm characteristics. Our results hold.

In other robustness tests, we use several alternative matching procedures. For example, we use propensity score matching to identify for each firm in the treatment sample the three closest matches from the subsample of non-California-headquartered firms.[17] Alternatively, we only choose the single closest propensity score matched firm, reducing the size of the control sample. We also conduct robustness tests in which, for each firm in the treatment sample, we select one non-California-headquartered firm that shares the same primary two-digit SIC code and is closest in terms of total assets. As in all matching procedures that we employ, we allow one specific control firm to serve as a matched control to more than one treatment firm but include any control firm only once in our sample. Finally, we also conduct robustness tests in which we use all non-California-headquartered firms as control sample. In general, results are very insensitive to the use of alternative control samples.[18]

4 The impact of the gender quota on firm performance

4.1 Main results

To analyze the market reaction to the introduction of the quota, we estimate market-adjusted firm value changes at firms subject to the quota and, in order to control for concurrent effects, compare them to market-adjusted firm value changes at firms not subject

[17] Propensity scores are estimated using the following explanatory variables: ROA, Leverage, ln(Total assets), PPE / AT, R&D/AT, and 2-digit SIC code industry dummies.

[18] We have also experimented with the addition of common support constraints by means of a restriction imposed for differences in total asset assets (+/- 50% of the California-headquartered firm) when using industry-size matching or by adding a caliper restriction (0.01) when using propensity score matching. Applying such restrictions leaves our results virtually unchanged.

to the quota. To this end, we regress different abnormal return measures on a treatment indicator, that is, a dummy variable set equal to one if a firm's headquarters are located in the state of California and zero otherwise. Standard errors are robust and clustered by two-digit SIC code level. The event date is Monday, October 1, 2018, the first trading day after the public announcement of the quota by the Governor's office. The announcement was made on Sunday, September 30, after the Governor had signed the law earlier that day. In our baseline setting, the sample consists of all California-headquartered firms that pass the filters described in Section 3.1 and three industry- and-size matched control firms per treatment firm as described in Section 3.3 (with replacement and avoiding double-counting). Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms.

Panel A of Table 3 reports the results from such difference-in-differences regressions using six different abnormal return measures, estimated over different event windows that range from one to five days in length. The results in Column 1 show that the abnormal return on the first trading day after the announcement (i.e., on October 1, 2018) is 0.45% lower for California-headquartered firms than for the matched control firms. The coefficient on the dummy variable whether the firm is headquartered in California is significant at the 1% level (t-statistic of -3.65). The results in Column 2 indicate that also on the second trading day after the announcement (October 2), California-headquartered firms significantly underperform the sample of matched control firms by -0.35%. When we compute cumulative abnormal returns as the sum of the daily returns over the two days following the announcement of the introduction of the gender quota, in Column 3, we find 0.73% lower returns for California-headquartered firms compared to the sample of matched control firms. The obtained estimate is not only economically relevant but also statistically significant with a t-statistic of -3.22. These results remain statistically significant and economically meaningful when we extend the event window to a symmetric three-day event window (Column 4), an asymmetric three-day event window (Column 5), or a symmetric five-day event window (Column 6). With -1.19%, the latter is comparable to Ahern and Dittmar (2012), who report a five-day abnormal stock returns of -3.54% for Norwegian firms with no female directors around the quota adoption, compared to -0.02% for firms with at least one female director. In California, the five-day announcement effect appears to be economically smaller than the effect in Norway. This may have to do with the fact that penalties in Norway are higher (i.e., the firm would be liquidated if it violates the quota requirement) and that there is still uncertainty in California whether the bill is legally valid.

The identifying assumption central to a causal interpretation of our difference-in-differences results in Table 3 is that treated and control firms share parallel trends before the onset of treatment, i.e., the introduction of the quota on September 30. Panel A of Figure 2 shows mean daily abnormal returns around the event date for the California-headquartered firms and for size- and industry-matched non-California-headquartered control firms. Panel B shows differences in mean daily abnormal returns between the two groups. In Panel A, the three pre-treatment observations (September 26 to September 28) confirm that the parallel trends assumption is fulfilled as the trends in pre-treatment abnormal returns are indistinguishable between the two groups of firms. On the event day, i.e., the first trading day after the announcement of the introduction of the quota (October 1), and the subsequent day (October 2), however, the abnormal returns of California-headquartered firms are significantly more negative by -0.45% and -0.35%, respectively (t-statistics of

-2.29 and -1.90).[19] On October 3, the difference in mean abnormal returns shrinks to an insignificant -0.15%. Panel B of Figure 2 confirms these patterns and again provides supportive evidence for the parallel trends assumption: There is clearly no difference in the pre-treatment trends across the two groups of firms.

In Panel B of Table 3, we test for the robustness of these results with respect to the model used to predict daily expected returns. Instead of applying a regression-based market model to predict expected returns in the event window as in Panel A, we rely on a market-adjusted model. That is, we subtract the daily market return from the observed daily stock returns to obtain daily abnormal returns. The market-adjusted model mitigates concerns of misestimated betas and simply assumes that each firm has a beta of one. The results remain qualitatively unchanged, but the economic magnitude of the coefficient estimates is, if anything, slightly reduced. For instance, the average two-day cumulative abnormal return difference between firms headquartered in California and the control firms is now -0.72% (Column 3) versus -0.73% in Column 3 of Panel A, but still significant at the 1% level (t-statistic of -3.39).

In Panels C and D, we vary the control group. In Panel C, we use propensity score matching to identify matched control firms. Specifically, we use profitability (ROA), leverage, firm size, asset structure (PPE / AT), R&D intensity (R&D / AT), and 2-digit SIC code industry dummies to estimate propensity scores. We then choose the three closest propensity score matched control firms. As with the size- and industry-matching applied in our baseline setting, a non-California-headquartered firm may serve as a matched control firm to more than one California-headquartered firm, but every control firm is included

[19] Results from tests for differences in means and tests whether means are significantly different from zero are reported in Table OA.2 in the Online Appendix. The mean abnormal returns of both the group of California-headquartered firms and the group of matched control firms are significantly negative on the first two trading days after the introduction of the quota (-2.46% and -1.73%; t-statistics of -11.56 and -10.10, respectively), suggesting that there may be spillover effects to non-California-headquartered firms or an anticipation of other states introducing quotas as well. We analyze potential spillover effects and channels through which they operate in Section 6 of the paper.

only once in our control sample. In Panel D, we use all non-California-headquartered firms that pass our sample selection procedure explained in Section 3.1 as control group. The results obtained in both panels are very similar to those reported in Panel A.

Panel E reports results from a set of additional robustness tests. For the sake of brevity, we only report results from tests using the two-day CAR (CAR(0,1)) as dependent variable. In Columns 1 and 2, we only select the single closest matched control firms in the size- and industry-matching procedure of Panel A and the propensity score matching approach used in Panel C, respectively. The coefficients on the dummy variable – firm is headquartered in California – are again similar to previous estimates, and, despite substantial reductions in sample size by about one third to 809 and 801 firms, respectively, remain significant at the 10% level or higher. The remaining four tests aim to mitigate concerns that a small number of observations in our sample may drive the results documented so far. The control sample is obtained using the same size- and industry-matching procedure as in Panels A and B. In Column 3, we test whether penny stocks drive our results and exclude all stocks in our sample with day-end closing prices below USD 1 on the event date from the sample (e.g., Amihud and Stoyanov, 2017). In Columns 4 and 5, we winsorize the cumulative abnormal returns at the 0.5th and 99.5th and 5th and 95th percentiles, respectively, instead of the 1st and 99th percentiles. Finally, in Column 6, we estimate median regressions without winsorization of the dependent variable instead of OLS regressions with winsorization to mitigate the effect of outliers. In all these tests reported in Panel E, we find the results to remain qualitatively similar.

Taken together, the results in Table 3 show that the adoption of the California mandatory gender quota law is associated with negative and significant announcement returns, suggesting that the introduction of the quota imposes significant costs on California-headquartered firms and their shareholders. Economically, the estimates vary but indicate that California-headquartered firms underperform non-California-headquartered control firms by between -0.51% and -0.99% over a two-day period that includes the two trading days after the announcement. As one would expect, the bulk of the abnormal market reaction occurred on the first trading day after the announcement. The result that California-headquartered firms lost substantial value following the quota's introduction is remarkably robust and remains if we use alternative event window lengths, construct different control groups, or take various measures to mitigate the impact of outliers.

4.2 Controlling for firm and industry characteristics

In Columns 1 to 4 of Table 4, we test whether our estimate for the loss in shareholder value around the law's announcement is robust to controlling for a set of firm-level covariates as well as industry-level effects. If treatment assignment is exogenous, and thus if our difference-in-differences analysis is valid, the inclusion of covariates should not materially affect our results obtained in Table 3 and discussed above. In Column 1 of Table 4, we therefore add a set of firm specific control variables to our baseline specification reported in Panel A of Table 3, Column 3. These variables capture a firm's return on assets, its leverage, its size, the fraction of total assets invested in property, plant and equipment, and the fraction of total assets invested in R&D expenditures. In Column 2, we additionally add two-digit SIC code industry dummies. These industry fixed effects absorb unobservable industry characteristics, such as industry shocks occurring simultaneously with the announcement of the quota adoption. In Columns 3 and 4, we vary the industry definition underlying the industry dummies: one-digit SIC industries (Column 3) and three-digit SIC industries (Column 4). Across all these specifications, the coefficient on the dummy indicating whether a firm is headquartered in California remains economically similar to Table 3 and is statistically significant at the 5% level.

4.3 Alternative event dates

Our event study results so far focus on the date on which the Governor of California signed the quota law. As Eckbo, Nygaard and Thorburn (2018) point out, it is important to include all major quota-related news events that increase the likelihood of a quota law in an analysis of changes in firm value. Therefore, we conduct a similar difference-in-differences analysis as described in Section 4.2 but vary the event dates. Specifically, in Columns 5 and 6 of Table 4, we estimate cumulative abnormal returns over a two-day event window that includes the quota law's introductory date (January 3) and the day after (January 4). In Columns 7 and 8, we estimate cumulative abnormal returns over a two-day event window that includes the successful Senate vote date (May 31) and the day after (June 1). Finally, in Columns 9 and 10, we use a three-day event window that includes the days of the successful Assembly vote and the second Senate vote, which took place on consecutive days (August 29 and 30), and the day after (August 31). All six coefficients on the California-headquarter indicator variable are positive and statistically insignificant at conventional levels, suggesting that the market reaction to the California gender quota was confined to the days after the Governor signed the law.

To further justify the choice of our event window, in Figure 3, we display the distribution of the newspaper coverage if we run article searches using Factiva for two different search terms, "California female board quota" (black line) and "California Senate Bill 826" (gray line). Panel A displays the weekly distribution of articles that contain these search words during the time period from December 1, 2017, to November 30, 2018. The figure shows that newspaper coverage of the quota, using either search word, is clearly concentrated in the week following the signature of the bill by Governor Brown on Sunday, September 30. Moreover, a reading of the articles published before September 30 confirms a considerable uncertainty whether the Governor will sign the controversial bill, as described in more detail in Section 2. Panel B of Figure 3 shows the daily distribution of articles in the period from August 1, 2018, to November 30, 2018. As expected, newspaper coverage peaks on October 1, the day after the bill was signed into law, supporting the use of a two-day event window that covers the first two trading days after the event.

4.4 Standard difference-in-differences estimates

In this section, we conduct a standard difference-in-differences analysis using a treatment dummy and a post-treatment dummy. To this end, we estimate OLS regressions of daily abnormal returns (ARs) on a dummy variable which is equal to one if a firm is headquartered in California (CA HQ (d)) and zero otherwise, and a dummy variable which is equal to one for observations measured after the implementation of the quota (Post (d)) and zero for observations measured before the implementation of the quota. We also add an interaction term between these two variables. Standard errors are clustered by firm-level. The results are reported in Table 5. Column 1 reports results obtained when using a four-day event window with two pre-treatment and two post-treatment observations per sample firm. Consistent with results reported in Tables 3 and 4, the difference-in-differences estimator, i.e., the coefficient on CA HQ (d) $\times$ Post (d), is negative and significant at the 5% level. In terms of economic magnitude, the coefficient suggests a two-day abnormal return of California-headquartered firms that is 0.76% lower than that of non-California-headquartered firms (0.38% per post-treatment day), a number reasonably close to the two-day abnormal return difference of between -0.51% and -0.99% reported in Table 3. The results in Columns 2 and 3 show that the inclusion of industry dummies and a set of firm-level control variables and industry dummies, respectively, leaves the results virtually unchanged. In Column 4, we add firm fixed effects to the specification in Column 1 to control for unobservable heterogeneity at the firm-level that is time-invariant. Note that the firm fixed effects absorb all firm-level covariates, including the treatment dummy, CA

HQ (d), as these variables are time-invariant over the sample period used in this analysis. While the t-statistic is slightly reduced, the difference-in-differences estimate remains economically unchanged, still suggesting a two-day abnormal return difference of 0.76%.[20]

4.5 Calendar-time portfolio analysis

In this section, we conduct robustness tests that address concerns arising from the fact that we study the market reaction of firms to one single event date. More specifically, all firms in the treatment group, i.e., the firms headquartered in California in our sample, are treated at the same date, October 1, 2018. Such a single event date may result in contemporaneous cross-correlation of (abnormal) stock returns. To address this concern, we follow Eckbo, Nygaard, and Thorburn (2018) and form equally-weighted calendar time portfolios of all California-headquartered firms, our treatment sample, and all non-California headquartered size- and industry-matched control firms, our control sample. We then estimate the portfolios' daily abnormal returns by estimating the following time-series regression over a sample that includes all observations from the 250-day estimation window, which ends on September 21, and the observations from the respective event window:

$$r_t = \alpha + ARd_t + \beta r_{wt} + \epsilon_t \tag{1}$$

where r_t is the daily equally-weighted return of the portfolio of all California-headquartered (or size- and industry matched control) firms in excess of the daily 1-month U.S. treasury bill rate. Alternatively, to analyze differences in abnormal returns between California-headquartered and non-California-headquartered firms, we define r_t as the daily difference in portfolio returns of California-headquartered and non-California-headquartered firms. r_{wt} is the daily value-weighted market index return in excess of the daily 1-month U.S.

[20] In unreported results, we extend the event window to a six-day period around the announcement of the quota. The difference-in-differences estimator remains statistically significant at the 10% level and economically comparable to our main results.

treasury bill rate. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. d_t is a dummy variable that takes a value of one for each day in the event window and zero otherwise. AR is the average daily abnormal portfolio return over the event days. Hence, estimates for the cumulative abnormal returns are obtained by multiplying the obtained coefficient for AR by the number of days in the event window. For instance, to obtain the two-day CAR(0,1), CAR(0,1), d_t takes a value of one in the two-day event window that includes the first two trading days after the quota came into effect. The two-day cumulative abnormal return, CAR(0,1), is then computed as 2 $\times$ AR.

The results are reported in Table 6. As in Table 3, we use six alternative abnormal return measures, estimated over event windows that range from one to five days in length. Consistent with results reported in previous tables and in Figure 2, we find announcement returns to the introduction of the quota to be significantly more negative for California-headquartered firms, as shown in the last two columns of the table. In terms of economic magnitude, the results obtained here are similar to those reported in Table 3. For instance, the two-day cumulative abnormal return, which includes the event day and the day after (CAR(0,1)), is -0.73% in Panel A of Table 3 and -0.74% in Table 6. Hence, accounting for potential contemporaneous cross-correlation resulting from a single event does not materially affect our results.

5 Cross-sectional tests

In this section, we analyze whether all California-headquartered firms are equally affected by the passage of the law or whether there are important cross-sectional differences. For instance, if the quota's introduction is expected to result in supply side frictions on the (female) director labor market, we would expect California-headquartered firms to be more (negatively) affected the fewer female directors they have on the board at the time of the introduction of the quota and the more directors are required to comply with the quota's requirements. Similarly, we would expect larger firms to be less (negatively) affected by the quota as larger firms are generally more attractive employers to directors as both compensation and reputation are strongly positively correlated with firm size.

Following Ahern and Dittmar (2012), we analyze such cross-sectional patterns using pooled OLS regressions of two-day cumulative abnormal returns (CARs) on a dummy variable that is equal to one if a firm is headquartered in California (CA HQ (d)), a firm-specific characteristic, e.g., related to quota compliance or female board representation, and an interaction term between the dummy whether the firm is headquartered in California and the firm characteristic. In other words, we use the regression framework from Column 3 in Panel A of Table 3, and saturate it with a firm-characteristic and an interaction term between a firm-characteristic and the California-headquarter indicator. Results are reported in Tables 7, 8, and 9. In the first column of Table 7, we interact the California-headquarter dummy with a variable equal to one if a firm is, as of September 30, not complying with the quota, and zero otherwise. The results imply that California-headquartered firms not in compliance with the quota underperform California-headquartered firms in compliance with the quota by -1.03% (significant at the 1% level). The coefficient on the stand-alone variable indicating whether a firm complies with the quota is insignificant. The stand-alone coefficient on the California-headquarter indicator is positive and significant at the 5% level, suggesting that California-headquartered firms, that elected female directors early and were in compliance before the introduction of the quota, may even benefit from the introduction of the quota. Next, we analyze whether the market reaction at California-headquartered firms not only depends on whether firms fail to comply with the quota but whether it also depends on the extent to which the quota is not fulfilled. To this end, in Column 2, we use a variable that counts the number of female directors necessary to fulfill the quota (# female directors missing), and, in Column 3, we use this variable scaled by board size (Shortfall (%)). Both variables, when interacted with the California headquarter indicator, load significantly negative, suggesting that the market reaction indeed varies with the female director gap on the board of California-headquartered firms. Moreover, the coefficient on the stand-alone California-headquarter indicator variable is indistinguishable from zero once we control for the need of California-headquartered firms to appoint female directors, supporting the notion that stock return differences between firms headquartered in California and their matched peers around the introduction of the quota are not driven by other factors than the female director gap. Rather, these results imply that frictions associated with appointing female directors increase with the number of female directors needed to fulfill the quota.

In Column 4 of Table 7, we test whether the stock market response depends on the percentage of female directors on the board. The results show that California-headquartered firms show a more negative announcement return to the introduction of the quota in general – the standalone coefficient on CA HQ (d) is negative and significant at the 1% level – but that the announcement return becomes more positive the larger the percentage of female directors on the board. Indeed, the interaction term between the California-headquarter indicator and the percentage of female directors on the board is positive and significant. This finding is consistent with the results in Columns 1 to 3.

Besides using this setting to explain variation in abnormal stock returns around the quota announcement with board structure variables, we also aim to shed light on the question which firms face larger difficulties in appointing female directors. Both, anecdotal and survey evidence, suggests that director referrals play an important role in filling board vacancies. PwC's 2016 Corporate Director Survey, for instance, reveals that 87% of surveyed directors rely on recommendations from their own board's network when hiring new directors. In fact, director referrals seem to be more important than search firm recommendations (60%), management recommendations (52%), or investor recommendations (18%).

Consistently, Fahlenbrach, Kim, and Low (2018) show that directors are more likely to obtain additional directorships, if their current board is well-connected.

We thus test whether a board with a larger network of female directors mitigates the firm's negative announcement returns observed around the quota adoption. To this end, we use BoardEx data and count for each sample firm, as of September 30, 2018, the number of distinct female directors that have a connection to at least one board member of the sample firm and sit on at least one board of another listed firm. A female director of another listed firm is defined to have a connection to a board member, if she currently shares or shared in the past an overlapping work engagement at a firm other than the sample firm, graduated from the same university within one year, or is (was) active in the same social organization. Our first measure of the female network size is the logarithm of one plus this female network size variable. Results reported in Column 1 of Table 8 show that California-headquartered firms with more female directors in the board's network have significantly higher, i.e., less negative, announcement returns. Columns 2 to 4 show variations of this analysis. In Column 2, we restrict the female board connection variable to employment connections only. In Columns 3 and 4, we address the concern that the size of the female director network may capture the size of the board's overall network by scaling the two variables used in Columns 1 and 2 by the total number of directors in the board's network. Results are similar across these variations. In summary, results in Table 8 suggest that more board connections to female directors mute the negative influence of the female board quota. Remarkably, we find a negative and statistically significant coefficient on the female network proxy variable across all four columns. One explanation for this result may be that California-headquartered firms with higher network connections are more likely to absorb female directors from the director labor market via such network connections. Hence, non-California-headquartered firms may lose valuable female directors.

Next, we study the role of firm size by interacting the California-headquarter indicator with proxies for firm size. As directorships at larger firms are associated with both higher compensation and higher reputational gains, we would expect larger firms to face fewer difficulties in attracting skilled female directors. Moreover, larger firms hire directors internationally, which likely gives access to a larger pool of female director candidates. In Column 1 of Table 9, we interact the California-headquarter dummy with the logarithm of total assets as a continuous measure of firm size, and, in Column 2, with a dummy variable that is equal to one for firms in the highest size quintile. The results in both columns suggest that firm size indeed mutes the negative quota effect, which means that in particular small firms with headquarters in California are responsible for the negative abnormal stock price reaction at California-headquartered firms observed around the adoption of the quota announcement. In particular, the results in Column 2 suggest that the negative stock return effect to California-headquartered firms is confined to firms in the smaller four size quintiles, but not to firms in the largest size quintile. Indeed, the coefficient on the interaction term is positive and slightly larger than the coefficient on the dummy variable whether the firm is headquartered in California (the combined effect of 0.19 is insignificant with a t-statistic of 0.5). Moreover, firms in the largest size quintile more generally do not experience negative announcement returns that are significantly different from zero: The coefficient on the largest size quintile dummy equals 1.68 and the intercept is -1.92, both significant at the 1% level, resulting in an insignificant combined effect of -0.24. Consistent with the results in Columns 1 and 2 of Table 9, Panel B of Figure 1 shows that larger firms (as proxied by index membership) have higher female board representation already before the quota's introduction.[21]

Finally, in Column 3 of Table 9, we study the role that corporate governance plays for the abnormal stock returns following the quota law announcement. To this end, we match

[21] One concern with these results is that larger firms have higher female board representation at the onset of treatment (i.e., the announcement of the quota's introduction). However, we continue to find a statistically significant size effect once we control for female board representation.

corporate governance data from Factset's SharkRepellent database, which enables us to construct, for each sample firm as of the quota adoption date, a corporate governance index in the spirit of Bebchuk, Cohen, and Ferrell's (2009) E-Index. Specifically, we construct a score variable to which we add one point for each of the following provisions: a poison pill, a staggered board, a supermajority vote requirement for mergers, a supermajority vote requirement to amend the charter, or a supermajority vote requirement to amend the bylaws.[22] The results show that the coefficient on the interaction term between the California-headquarter indicator and the (modified) E-Index is negative and significant, implying that California-headquartered firms with poor governance (i.e., more anti-takeover protection devices in place) suffer lower returns around the adoption announcement. These results suggest that firms' corporate governance structures or proxies of firm governance may influence firms' ability to attract (or willingness to appoint) qualified female directors.

The documented decline in the market value of equity of California-headquartered firms around the quota adoption announcement is economically relevant. Following the reasoning in Ahern and Dittmar (2012), a potential source of this value loss is a deterioration of board quality induced by the quota which restricts the appointment of new directors. An alternative explanation for such a value loss would be that California-headquartered firms were negatively affected by the quota's introduction because it shows California's willingness to legislate non-economic values on California firms. If this were true, the value loss is not about female board representation, but about other corporate actions. To test this, we use the KLD Stats database and construct a sustainability index in the spirit of Servaes and Tamayo (2013). KLD categorizes corporate social responsibility-related items into seven

[22] We do not use ISS data, which is the data source of the E-Index in most papers. The clear advantages of SharkRepellent are that the data are already available for September-end 2018 and that they cover firms outside of the S&P 1500 index. The drawback is that SharkRepellent does not provide information on one of six provisions of the E-Index, that is, whether the CEO's contract contains a severance agreement provision known as a "golden parachute". Hence, our modified E-Index is available for most sample firms but only includes five attributes, ranging from zero to five with a higher score indicating weaker shareholder rights.

different categories: “Community”, “Corporate Governance”, “Diversity”, “Employee Relations”, “Environment”, “Human Rights”, and “Product”. For the purpose of this test, we omit the “Corporate Governance” category, because prior research shows that KLD’s assessment of governance quality differs significantly from that employed in the finance literature (e.g., Krüger, 2015). Hence, we resort to the E-Index as a proxy for corporate governance, as explained above. We also exclude the “Diversity” category because it is closely related to and might thus capture female board representation. For each category, KLD defines a set of binary indicator variables, which are either positive (“Strengths”), or negative (“Concerns”). As the number of strengths and concerns varies across categories, we divide the number of strengths (concerns) for each firm-year within each category by the maximum possible number of strengths (concerns) in each category-year to obtain two indices that range from zero to one. Within each category, in each firm-year, we subtract the concerns index from the strengths index to end up with a net sustainability score that ranges from –1 to +1. Finally, we sum up the five different category indices to obtain our sustainability score that ranges from -5 to +5 with a higher score indicating a more sustainable firm. As KLD regularly changes the list of strengths and weaknesses within the categories, which results in a large time-series variation of the measure, we use the mean sustainability score of all available observations on each sample firm, a maximum of seven yearly observations (2010 – 2016).[23] The results on the sustainability index are reported in Column 4 of Table 9. They show a positive and significant coefficient on the interaction term between the California-headquarter dummy and the sustainability index as well as on the stand-alone sustainability index. The coefficient on the California-headquarter dummy remains negative and significant. Hence, the market reaction to the quota is more positive (or less negative) the higher the sustainability score, both at California-headquartered firms and at control firms. Taken together, these results imply that some but not all of

[23] Note that data coverage in the KLD Stats database currently ends in 2016.

the market reaction around the quota announcement is driven by California's willingness to legislate non-economic values on California firms.

6 Spillover effects

In this section, we analyze potential spillover effects of the gender quota's introduction to firms headquartered in other states than California. First, we analyze whether non-California-headquartered firms experience a negative and significant stock market reaction to the introduction of the quota as well. Second, we test for the channels through which these spillover effects materialize.

Results reported in Panel A of Figure 2 and Table 6 indeed show negative and significant stock market reactions to the introduction of the quota not only to California-headquartered, but also to non-California-headquartered firms. In Table OA.2, we conduct our baseline event study analysis, underlying the analysis in Panel A of Table 3, for the subsamples of California-headquartered and industry- and size matched control firms separately. The results show a significantly negative stock market reaction of -1.73% (t-statistic of -10.10) for non-California headquartered control firms over the first two trading days after the introduction of the quota in California. In summary, standard event study analysis (Table OA.2) and a calendar time portfolio analysis (Table 6) are suggestive of spillover effects of the quota law to non-California-headquartered firms and indicate that these firms, even though not directly affected by the law, still show a significantly negative market reaction to the announcement of the quota.[24]

There are mostly two channels through which such spillover effects are expected to

[24] The existence of such spillover effects violates the stable unit treatment value assumption (SUTVA) underlying any difference-in-differences approach. Specifically, this assumption postulates that the outcome of some unit of observation should be unaffected by the assignment of treatment to other units. Given that the control firms also experience negative stock returns, these spillover effects most likely lead to an underestimation of the firm value effect of the quota law, that is, the return differences between California-headquartered firms and the sample of control firms.

operate. First, California-headquartered firms may attempt to attract female directors from the boards of non-California-headquartered companies to be able to comply with the quota. Hence, competition among firms for skilled female directors may become more intense, and some firms may even be expected to lose valuable female directors. Second, if the introduction of a mandatory gender quota in California raises concerns that other states follow by also introducing gender quotas, we would expect firms headquartered in states that are more likely to pass such laws to show more negative announcement returns to the quota. To empirically test these two channels, we retain only the matched non-California-headquartered firms in our sample and estimate regressions similar to those reported in Tables 7 to 9. Specifically, we regress the cumulative abnormal return over the two trading days after the announcement of the quota on a firm-specific characteristic, e.g., female board representation.

To test for the existence of the first channel, we need a measure that quantifies the extent to which an increase in the demand for female directors at California-headquartered firms hampers the ability of non-California-headquartered firms to retain or appoint female directors. We construct two measures, which both rest on the assumption that similar firms compete for the same directors and that director supply is limited (e.g., Knyazeva, Knyazeva, and Masulis, 2013). To construct the first measure, we count, for each matched control sample firm, the total number of female directors required for all California-headquartered firms in the same two-digit SIC code industry to comply with California's female director quota (# female directors missing). To address the problem that this measure is correlated with the number of firms in an industry, we scale it by the total number of female directors required by all firms in a two-digit SIC code industry to comply with the Californian quota, i.e., assuming that all firms – whether headquartered in California or elsewhere – have to comply to such a quota to obtain our second measure. In Column 1 of Table 10, where we use the first of the two measures, we obtain a negative coefficient that is significant at the 10% level. In Column 2, we use the scaled variable and find the coefficient to be negative and significant at the 1% level. These results confirm that non-California-headquartered firms in industries that likely face stronger competition for female talent in the boardroom experience more negative announcement returns in response to the introduction of the quota.

Results in Tables 7 and 9 show that the market reaction to the quota's announcement was less negative for California-headquartered firms the higher the fraction of female outside directors on the board and the larger the firm. In Column 3 of Table 10, we test whether we find similar patterns in the spillover effects to non-California-headquartered firms. To this end, we interact a dummy variable indicating whether a firm is in the highest total assets quintile with the fraction of female directors on the board while controlling for the two stand-alone variables. The results show that a larger female board representation is associated with more negative spillover effects. However, the coefficient on the interaction term is positive, significant, and larger in magnitude than the coefficient on the female board representation variable, suggesting that the negative spillover effect is not only muted for large firms but even reversed. Hence, smaller firms not headquartered in California may be at risk of losing valuable female directors to (larger) California-headquartered firms. In Column 4, we follow the idea that certain directors at non-California-headquartered firms may receive offers to join the boards of California-headquartered firms. As a proxy for the propensity to leave a given firm, we compute the natural logarithm of the average tenure of the female directors on the board. Naturally, this measure can only be computed for firms with at least one female director on the board, so the number of observations is reduced. Still, we find that the tenure of the female directors on the board is positively correlated with the returns of non-California-headquartered firms, consistent with the notion that firms experience a more negative stock price reaction the more likely it is to lose a female director.

In Table 11, we explore the second channel: The negative spillover effect might also be driven by the propensity of a given non-California-headquartered firm to become subject to similar gender quota regulation in the future. Given the opposing views of the Democratic and Republican parties on the controversial issue of gender legislation for corporate boards, we use the local political orientation as a proxy for the propensity to become subject to a similar quota. To this end, we collect the state-level results of the 2016 Presidential Election for the Republican and Democratic Party and assign firms to states based on the location of their headquarters. In Column 1 (2), we find that firms with headquarters outside California but in states with a higher share of votes obtained by the Democratic (Republican) party experience more negative (positive) announcement returns around the adoption of the gender quota in California. Similarly, in Column 3, we use an indicator variable labeled Trump (d) which is equal to one if a firm is headquartered in a state where the majority of the votes in the last presidential election were obtained by the Republican Party. Consistent with the results shown in the previous two columns, the coefficient on this variable is positive and significant.[25]

After California's introduction of non-binding board gender legislation in 2013, five states adopted similar legislation. These states are Massachusetts, Illinois, Pennsylvania, Ohio, and Colorado. As such non-binding board gender legislation may be related to the probability of introducing a mandatory quota, we test whether the existence of such "soft quotas" is associated with more negative announcement returns around the introduction of the quota law in California. However, such a negative effect on announcement returns may be reduced, or even offset, if such soft quotas induce affected firms to prepare for

[25] Alternatively, we construct a variable that is based on the political orientation of the state-level legislature. To this end, we define an indicator variable which we set equal to one if, as of the quota adoption date, the Governor, the upper, and the lower house in the state in which a firm's headquarter is located are controlled by the Republican party. Unreported results show that the coefficient on this dummy variable is 0.74 and thus larger than the coefficient on the Trump dummy in Column 3, but not quite statistically significant at conventional levels (t-statistic of 1.62).

the introduction of a mandatory quota, for example by appointing female directors to their boards. In Column 4, we report results from a regression, which includes a dummy for the existence of a soft quota in the firms' headquarter state. The coefficient on this dummy is negative but insignificant, suggesting that returns between firms headquartered in soft-quota states and those headquartered in other states do not differ.

In order to further test the propensity of non-California-headquartered firm to become subject to similar gender quota regulation in the future, we examine cannabis legislation. Legalization of medical use of cannabis has been pioneered by California, and other states have followed suit. California was the first state to legalize the use of cannabis for medical purposes. This occurred in 1996 when California voters approved Proposition 215 – the Compassionate Use Act of 1996. The use of cannabis for medicinal purposes is now legal in 34 states and the District of Columbia. In 2016, California voters approved the Adult Use of Marijuana Act through Proposition 64. Ten more jurisdictions have legalized cannabis for recreational use (Alaska, Colorado, the D.C., Maine, Massachusetts, Michigan, Nevada, Oregon, Vermont, and Washington). We thus theorize that examining cannabis adoption rates in California's path and in accord with California will reflect a legislature's tendency to adopt a similar gender quota law as California.

We define three variables related to the states' policy on cannabis use. The first variable (Medical cannabis legal (d)) is a dummy variable that is equal to one if the firm is headquartered in a state that has legalized cannabis for medical use.[26] The second variable, Cannabis legalization (score), is a count variable that is equal to zero if any type of cannabis use is considered illegal, equal to one if the use of cannabis is legal for medical purposes, equal to two if the recreational consumption of cannabis is illegal, but has been

[26] States that have only legalized the use of Cannabidiol (CBD) oils are coded as zero because some of these states banned the active chemical ingredient, Tetrahydrocannabinol (THC) (e.g., Kansas) or set THC limits close to zero (e.g., Indiana), resulting in legal uncertainty amongst sellers and significantly reducing the availability of CBD oils as they can contain traces of THC.

decriminalized, and equal to three if the recreational consumption of cannabis is legal. Hence, this variable measures the extent to which state law contradicts federal law, under which cannabis is treated as a Schedule 1 drug, which prohibits use also for medical applications. The third variable is Quick medical cannabis follower (d). This is a dummy variable, which is equal to one if the medical use of cannabis has been legalized quickly (i.e., below the sample median of 16.5 years after California's legalization of medical use of cannabis). In tests using this variable, sample size is reduced because we only consider firms headquartered in states which have legalized the use of cannabis for medical treatment. The results are reported in Columns 5 to 7. The coefficients on all three variables related to the legal status of cannabis in the firms' headquarter state are negative, and two of them are marginally statistically significant, suggesting that firms headquartered in states that are more likely to follow California in the adoption of gender board quotas have significantly lower returns.

In summary, results in this section show that the introduction of the gender legislation in California was associated with negative announcement returns not only at California-headquartered, but also at non-California-headquartered firms. These findings are consistent with spillover effects of the California gender quota to non-California headquartered firms. Moreover, we find evidence in support of two channels through which these spillover effects may operate: First, firms that are more likely to face difficulties in appointing or retaining female directors have more negative returns, and, second, firms that are more likely to become subject to a future gender quota law also have more negative announcement returns.

7 The impact of the quota on board structure

7.1 The impact of the quota on female board representation

Gender quotas introduced in other countries have been shown to be an effective way to force firms to increase female board representation. In Norway, for instance, the share of female board representation has increased from around 10% in 2003, the year the Norwegian Parliament passed the quota law, to about 40% in 2008, the year the quota had to be fulfilled (Ahern and Dittmar, 2012). In the case of California's quota, however, it is a priori unclear whether firms will respond by restructuring their boards for at least three reasons. First, the new legislation prioritizes one gender over the other, thereby potentially conflicting with the U.S. Constitution, the California Constitution, and Civil Rights Law. Second, it may violate the internal affairs doctrine, which postulates that matters involving the board of directors are subject to the laws of the state of incorporation and not where the headquarters are located (Grundfest, 2018). Not surprisingly, at least 30 business groups already indicated their willingness to challenge the law.[27] Third, while the penalty for non-compliance in Norway is a forced liquidation of the company, non-compliance with the Californian quota is penalized with a fine that seems rather small at least from the perspective of the larger firms in the sample.

Even though the adoption of the quota in California took place only recently, we conduct a preliminary analysis with the goal of testing whether California-headquartered firms respond to the quota law, that is, whether they already started to increase the share of female directors on the board relative to a sample of matched control firms headquartered outside California. Figure 4 shows the change in mean female board representation relative to the introduction of the quota on September 30, 2018, in percentage points on a daily

[27] See http://ctweb.capitoltrack.com/public/search.aspx?t=bill&s=SB826&go=Search&session=17&id-=1dae9efb-651d-4a02-a05d-360ca7965b14.

basis from July 1 to December 31 using board data from BoardEx. The solid lines show these changes for the treatment sample (in black) and the industry- and size matched control sample (in grey). Consistent with the notion that California's quota effectively and timely affects female board representation, California-headquartered firms appear to start responding to the quota by increasing female board representation relative to the sample of control firms already two weeks after the adoption of the quota law and continue to appoint more female directors to their boards than firms in the control sample until the end of our sample period in December 2018. Moreover, California-headquartered firms in need of at least two female directors to comply with the quota's requirement (black dashed line) seem to respond stronger when compared to the other firms headquartered in California and the non-California-headquartered control firms that need at least two female directors to fulfill the quota (grey dashed line).

To analyze this pattern in a regression framework, we estimate, for each firm, the fraction of female directors on the board at the end of September, October, November, and December 2018, which results in a firm-month panel containing up to four monthly observations per firm. We then regress the fraction of female directors on the board on dummy variables for the month of observation, omitting September, and interaction terms between the California headquarter dummy and the month dummy variables. To account for time-invariant unobserved heterogeneity at the firm level, we add firm fixed effects. We cluster standard errors at the firm level. The coefficients on the month dummies indicate the percentage points by which female board representation has changed, on average, across all sample firms compared to the introduction of the quota at the end of September 2018. The coefficients on the interaction terms are the difference-in-differences estimators, that is, the average difference in the change of female board representation of Californian firms relative to the control firms in a given month. If the quota law already had a statistically significant impact on female board representation at firms headquartered in California, the difference-in-differences estimators would show a positive and significant coefficient.

Results in Column 1 of Table 12 show that female board representation at California-headquartered firms indeed increased relative to the sample of non-California-headquartered control firms. Specifically, one month after the quota's introduction, the difference amounts to 0.19 percentage points and increases to 0.32 percentage points two months later. Yet these difference-in-differences estimates are statistically insignificant at conventional levels. As of December-end 2018, the difference increases to 0.45 percentage points, which is statistically significant at the 10% level. Compared to the average annual growth rate of female board representation, which amounts to 0.5% according to the 2018 report of the Corporate Women Directors International initiative[28], this monthly increase is economically meaningful.

Next, we test whether firms under more pressure to appoint female directors respond stronger to the introduction of the quota. In Column 2, we, therefore, retain only firms in the sample that need at least one female director to comply with the quota at announcement, and, in Column 3, we retain only firms that need at least two female directors. Consistent with our expectations and descriptive evidence provided in Figure 4, we find that the coefficients on the interaction terms between the California-headquarter dummy and the month dummy variables increase monotonically from Column 1 to Column 3. California-headquartered firms that require one (two) female directors to comply with the quota on average increased female board representation by 0.58 (0.69) percentage points relative to the control firms three months after the quota introduction. In the first three columns, the coefficients on all month-end dummy variables, which capture the general time trend, are positive and significant, suggesting that both California headquartered and non-California-headquartered firms significantly increased female board representation in the months after the introduction of the quota – but as the difference-in-differences esti-

[28] See https://globewomen.org/CWDINet/index.php/2018-fortune-global-200-companies/.

mates show, California-headquartered firms even more so.

The public debate around female board representation often emphasizes the number of firms without any female directors on the board to stress the most extreme cases of gender inequality.[29] The goal of our next test is to ascertain whether the new gender legislation has helped female directors to break into all-male boardrooms or whether the increase in female board representation documented in Columns 1 to 3 is mostly driven by firms that already have at least one female director on the board and now add additional female directors. Column 4 reports results obtained from estimating the regression in Column 1 and replacing the dependent variable with a dummy variable that is equal to one if a firm has no female director on the board at the end of a given month. The coefficients on all three difference-in-difference estimators are negative and statistically significant at the 10% level or higher. They suggest that the fraction of California-headquartered firms without a female director has gone down by roughly 3% relative to the matched control firms by December-end 2018. Hence, California's gender quota indeed seems to have induced some firms without any female director to appoint at least one. Moreover, these results indicate that California-headquartered firms move towards fulfilling the first threshold stipulated by SB 826, that is, that all firms headquartered in California have to have at least one female director by the end of the year 2019.

The next tests aim at dissecting whether the increase in female board representation by California-headquartered firms reported in Columns 1 to 3 is achieved by adding female inside or female outside directors. To this end, we re-estimate the regression in Column 1, but use the fraction of female outside directors (Column 5) and the fraction of female inside directors (Column 6), respectively, as dependent variables. The results in Column 5 closely mirror those in Column 1, while coefficients in Column 6 are all insignificant. Hence, firms

[29] See, for instance, Joanna S. Lublin, Why Breaking Into the Boardroom Is Harder for Women, The Wall Street Journal, Feb. 2, 2018.

under pressure to appoint female directors seem to primarily add female outside directors to their boards, at least in the first three months after the quota adoption.

If boards of California-headquartered firms are reluctant to ask male directors to resign or not propose them for reelection or if no potential female candidates with equal skill levels are available, California-headquartered firms might decide to comply with the quota by adding female directors without replacing any male directors. If this were a common response to the quota's introduction, we would expect to observe a significant increase in board size at California-headquartered firms relative to the sample of control firms. We test this conjecture by re-estimating the regression reported in Column 1 but use the logarithm of board size as the dependent variable. The results, reported in Column 7, provide some evidence that appointments to the boards of California-headquartered firms, in particular those shortly after the quota's introduction, resulted in an increase in board size. Indeed, all coefficients in this column are positive, but only the coefficient on the October-end difference-in-differences estimator is statistically significant.

In summary, these results imply that California's female board quota, although argued to be lacking teeth, is in fact taken seriously by firms: It triggered a significant increase in female board representation and a significant reduction in the number of firms without any female director on the board already within three months after adoption of the quota.

7.2 The impact of the quota on director skills

In this subsection, we test whether the female directors added to the board of California-headquartered firms after the introduction of the quota differ in terms of skills from incumbent directors and leaving directors. Such a shift would be expected if California-headquartered firms appointed male directors with a certain skill set before the introduction of the quota, and are now forced to dismiss some of these directors when recruiting female replacements from a pool of candidates with different characteristics. To test this empirically, we follow Ahern and Dittmar (2012) and compare the characteristics of newly appointed, leaving, and incumbent directors for the time period between the introduction of the quota (September 30, 2018) and the last month-end for which board data is available from BoardEx (December 31, 2018). In the following tests, we focus on California-headquartered firms and compare the characteristics of newly appointed female directors to the characteristics of incumbent female and male directors (Panel A of Table 13) and to the characteristics of leaving female and male directors (Panel B of Table 13). Table 13 shows that California-headquartered firms appointed 103 directors (46 or 45% of them females), while 115 directors (13 or 11% of them females) left the boards. The matched sample of incumbent directors in Panel A comprises all directors of firms in the appointment sample as of the quota adoption announcement. Note that we include the incumbent directors of a firm with multiple appointments only once in the control sample. For each director, we compute a range of measures that proxy for her characteristics and show results for a selected subset.[30]

Not surprisingly, results show that newly appointed female directors are younger than incumbent female and male directors as well as leaving female directors. More telling is a comparison of newly appointed female directors and newly appointed male directors in Panel B, which also indicates that newly appointed female directors tend to be younger, but the difference is not significant at conventional levels (t-statistic of -0.36). We also study more sophisticated measures of experience than director age. Results show that newly appointed female directors possess less work experience at another firm in the same two-digit SIC code industry (industry experience), less experience as a director of another listed firm, and less experience as an outside director of another listed firm than incumbent

[30] For newly appointed directors, we compute time-varying measures as of the date she joins the board of a given firm. For leaving directors, we compute time-varying measures as of the date she leaves the board of a given firm. For incumbent directors, we compute time-varying measures as of the quota adoption date (September 30, 2018).

female and male directors (Panel A). Although the majority of the results for comparisons between newly appointed and leaving female directors are similar in economic terms, they lack statistical power due to the small sample sizes. We also test whether newly appointed female outside directors differ from the other groups of outside directors in terms of their current occupation. These tests provide some evidence that newly appointed female outside directors are more likely to be employed in lower-level executive positions at other firms, for instance, as a Vice President, and are less likely to be employed as top executives at other firms, for example, as an inside director or a CEO, but differences are not statistically significant for all comparisons. Finally, we find some evidence that newly appointed female directors are more likely to be appointed as outside directors than as inside directors, supporting the finding from Section 7.1 than firms adhere to the quota by appointing female outside directors rather than female inside directors. Moreover, the appointed female outside directors are less likely to qualify as an independent outside director when compared to incumbent female and male outside directors.[31]

Even though these results are based on a small sample of board changes and should be considered preliminary, it is still important to emphasize that several of the characteristics underrepresented among female directors appointed to California-headquartered firms following the introduction of the quota have shown to be significantly related to firm performance. For instance, Drobetz, Oesch, Schmid, and von Meyerinck (2018) and Hoitash, Hoitash, and Faleye (2018) show that director industry experience is positively related to firm value. Fich (2005) and Fahlenbrach, Low, and Stulz (2010) find that directors simul-

[31] There is disagreement in the literature as to whether the independence data provided by BoardEx indeed reflects regulatory independence as provided, for example, by the ISS director dataset. Masulis and Zhang (2018), for instance, rely on BoardEx' role descriptions to classify directors as independent. In contrast, Fahlenbrach, Kim, and Low (2018) construct their own independence measure by classifying directors as not independent if they were employed by the focal firm in the past using BoardEx' employment data. Our own conversations with BoardEx' staff support the use of a director's role description to classify outside directors as independent, but we caution the interpretation of results related to director independence.

taneously serving as CEOs in other firms are associated with higher firm values. Knyazeva, Knyazeva, and Masulis (2013) document similar findings for independent outside directors.

In summary, these preliminary results are in line with both results reported in Ahern and Dittmar (2012) for the Norwegian quota and anecdotal evidence on female director characteristics in the US.[32] Moreover, these results are consistent with the notion that reductions in value at firms headquartered in California around the adoption of the quota are at least in part driven by shifts in investors' expectations regarding future board quality, likely caused by a limited supply of skilled female candidates that prevent firms from appointing directors that maximize firm value.

8 Conclusion

Even though the percentage of female university graduates increases each year and women successfully establish themselves in many academic profiles, such as economics, law, and medicine, women are still heavily underrepresented in corporate leadership positions. Many countries around the world recently introduced or currently discuss a board gender quota to increase female representation on corporate boards. In the U.S., California is the first state to introduce a mandatory board gender quota. Our results show significantly negative announcement returns for affected firms, and also negative spillover effects to stock returns of firms in other states that are likely to follow California's lead. Thus, they cast doubt on the view that quotas are an efficient way of breaking the glass ceiling for women in the corporate sector. Further, at least in the first three months after its introduction, the quota leads to increased hiring of less experienced female directors, which may generally harm the perception of women selected under quota programs. It remains to be seen,

[32] See Lucinda Shen, Some Female Leaders Aren't Huge Fans of California's New Law Requiring Women on Boards, Fortune, Oct. 4, 2018 (Tricia Griffith, one of the few female CEOs in an S&P500 firm: "We need to look deeper in organizations [for female candidates]. If you want someone that has a CEO role to their name, you're going to be very limited. So you need to go a layer deeper, then another layer. And those women and people are hard at work in operating roles that have a lot of value.").

however, whether the negative short-term reaction to the gender quota law is reversed in the future and whether firms ultimately benefit from the law, as was hoped for by its proponents.

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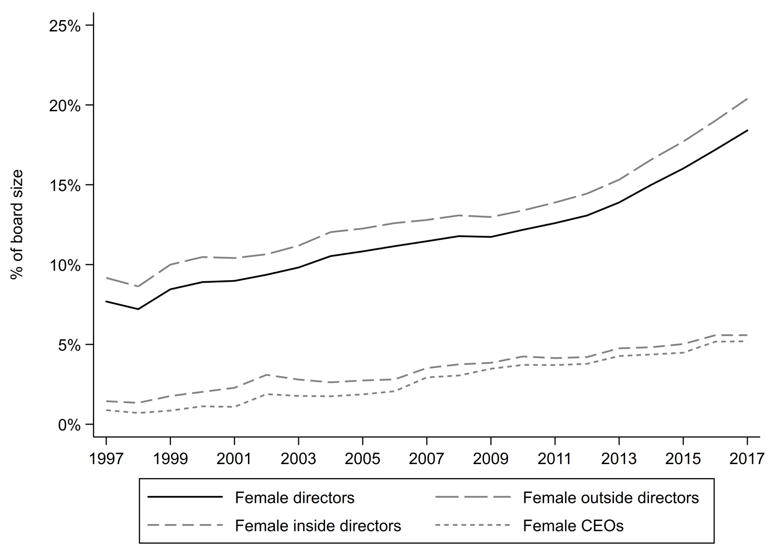
Figure 1: Female board representation in listed U.S. firms

This figure shows female board representation in all listed U.S. firms covered by the ISS (formerly RiskMetrics) Director database from calendar years 1997 to 2017. Panel A shows female board representation for different director types expressed as a fraction of board size. Inside directors are those directors classified as executives by ISS, CEOs are inside directors with current occupation description being "CEO", and outside directors are all directors that are not inside directors. Panel B shows female board representation by index membership. Board data is measured as of a firm-year's annual meeting date. Observations are assigned to the calendar year in which an annual meeting takes place. Index membership is determined as of the annual meeting date using index constituent data provided by Compustat.

Panel A: Female board representation by director type



Panel B: Female board representation by index membership

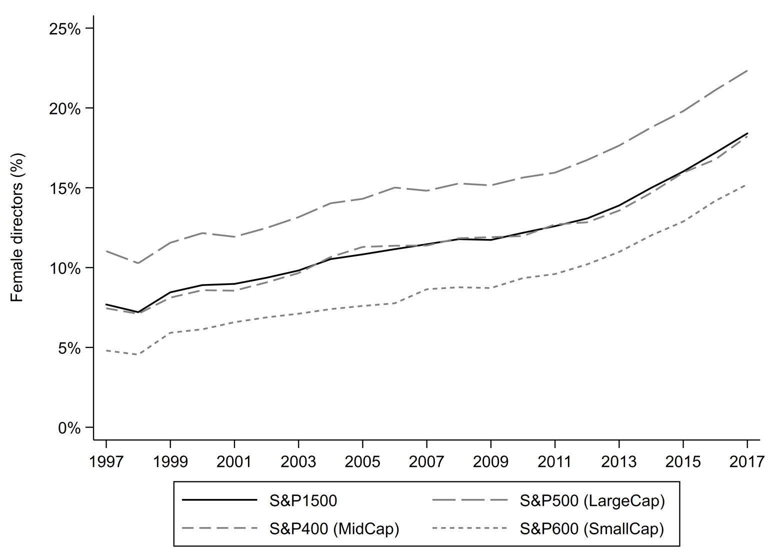
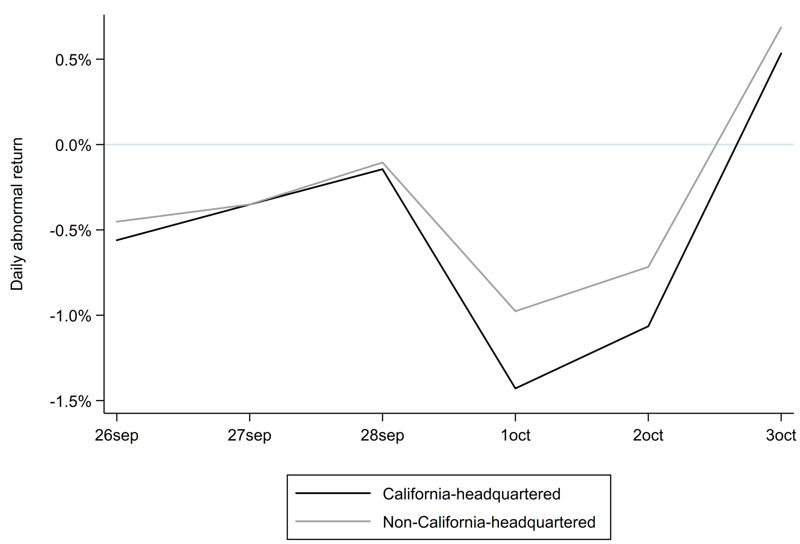


Figure 2: Abnormal returns around the event date (September 26 to October 3, 2018)

Panel A of the figure shows mean daily abnormal returns of 456 California-headquartered and 777 industry- and size-matched non-California-headquartered control firms. Panel B displays differences in daily abnormal returns between the sample of California-headquartered firms and the control sample of non-California-headquartered firms. Sunday, September 30, is the quota's adoption announcement date and October 1 the first trading day after the announcement, i.e., the event date. Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The sample comprises all firms in Compustat with a data entry within one calendar year before September-end 2018, excluding utility and financial firms (SIC codes 4940-4949 and 6000-6999, respectively), firms with missing information on the state in which it is headquartered, firms headquartered outside the US, firms with negative book value of equity, firms with missing financial control variables, firms that only list American Depository Receipts, and firms without a listing on NYSE, AMEX, or NASDAQ. We also require at least 125 daily return observations during the 250-day estimation window and complete return data for the entire five-day event window around the event date (October 1) and availability of board data from BoardEx. For each firm headquartered in California, we draw the three closest size-matched non-California-headquartered firms in the same two-digit SIC code industry as control firms. While the same firm may serve as a matched control firm to more than one California-headquartered firm, every control firm is included only once in our sample. All abnormal return measures are winsorized at their 1st and 99th percentile.

Panel A: Daily abnormal returns of California- and non-California-headquartered control firms



Panel B: Differences abnormal returns between California- and non-California-headquartered control firms

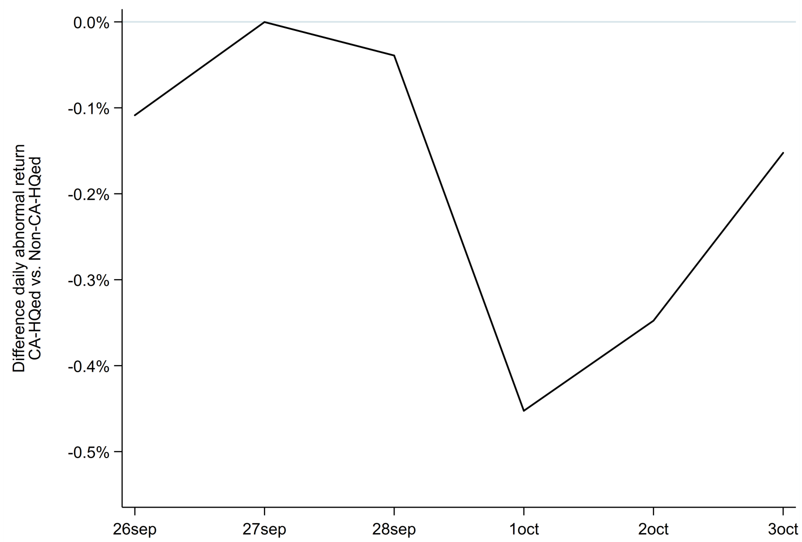
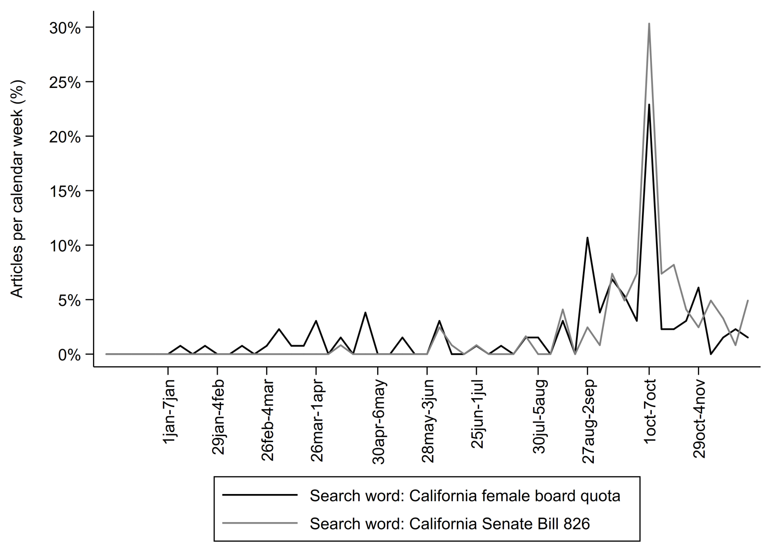


Figure 3: Distribution of newspaper coverage over time

This figure displays the distribution of newspaper coverage if we run article searches using Factiva for two different search terms, "California female board quota" (black line) and "California Senate Bill 826" (gray line), allowing for variations, e.g., "SB 826". Panel A displays the weekly fraction of articles that contain these search words during the time period from December 1, 2017, to November 30, 2018. Panel B shows the daily fraction of all articles published that contain these search words for the time period from August 1, 2018, to October 31, 2018.

Panel A: Weekly distribution of newspaper articles, December 1, 2017, to November 31, 2018



Panel B: Daily distribution of newspaper articles, August 1 to October 31, 2018

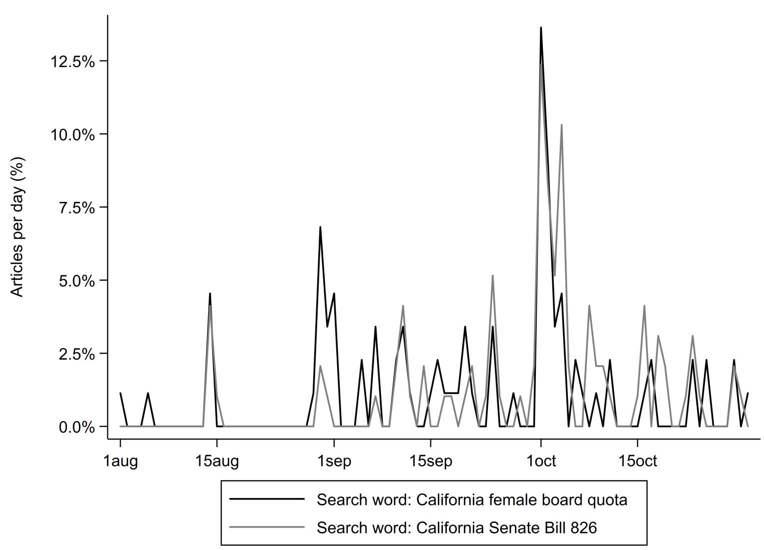


Figure 4: Changes in female board representation around quota adoption
This figure shows daily changes in female board representation for different firm samples relative to the introduction of the quota on September 30, 2018, over the time period July 1 to December 31, 2018. Solid lines show percentage point changes for a sample of California-headquartered firms (in black) and a sample of industry- and size-matched non-California-headquartered control firms (in grey). Dashed lines show changes in percentage points for subsamples of California-headquartered (in black) and non-California-headquartered firms (in grey) firms that miss more than one female director to fulfill California's board quota at the quota's introduction date. The sample comprises all firms in Compustat with a data entry within one calendar year before September-end 2018, excluding utility and financial firms (SIC codes 4940-4949 and 6000-6999, respectively), firms with missing information on the state in which it is headquartered, firms headquartered outside the US, firms with negative book value of equity, firms with missing financial control variables, firms that only list American Depository Receipts, and firms without a listing on NYSE, AMEX, or NASDAQ. We also require at least 125 daily return observations during the 250-day estimation window that ends September 21 and complete return data for the entire five-day event window around the event date (October 1) and availability of board data from BoardEx. For each firm headquartered in California, we draw the three closest size-matched non-California-headquartered firms in the same two-digit SIC code industry as control firms. While the same firm may serve as a matched control firm to more than one California-headquartered firm, every control firm is included only once in our sample. Board data is obtained from BoardEx. The red vertical line indicates the date the law was signed by the Governor (September 30, 2018).

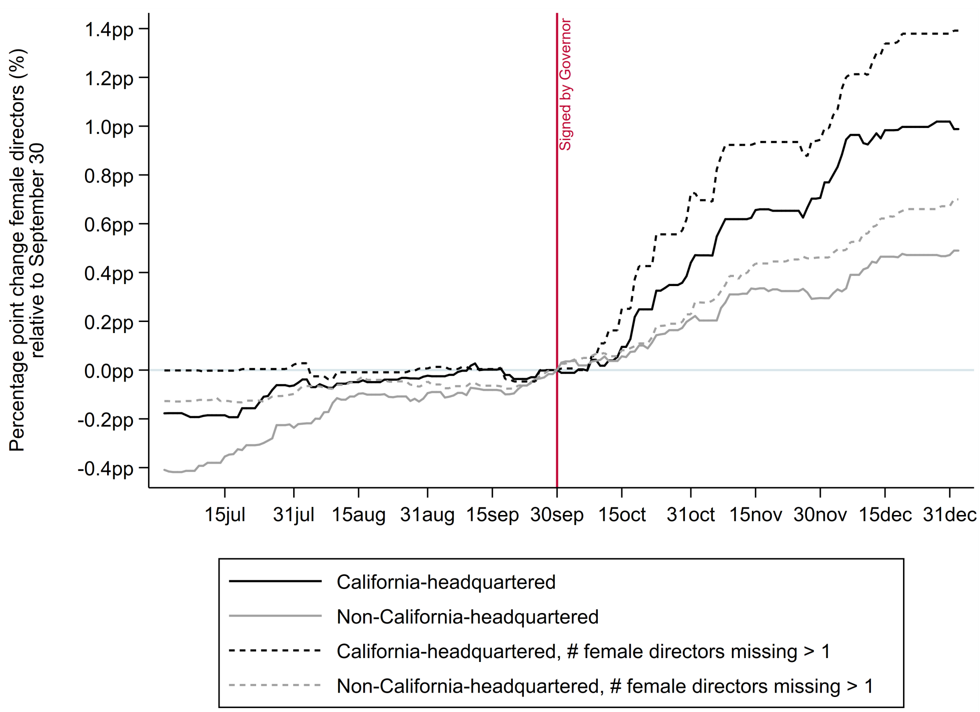


Table 1: Overview of international gender legislation

This table shows an overview of international gender legislation. We only include laws that apply to companies (not to state-owned entities), that are mandatory (not compulsory or on a comply-or-explain basis), and that carry some sort of pre-defined sanction (no potential sanctions). In the column labeled Focus, BoD indicates board of directors, SM senior management, and SB supervisory board (or "Aufsichtsrat") in Germany, which has a two-tier board system.

Country	Passage date	Applicable to	Focus	Requirement	Compliance by
Norway	19.12.2003 / 09.12.2005	Listed firms	BoD	1 each gender (2-3 directors), 2 (4-5), 3 (6-8), 4 (9), 40% (> 9)	2005 / 2008
Israel	11.03.2007	Listed firms	BoD	1 woman	2010
Iceland	04.03.2010	Listed and private firms	BoD	40% each gender; if < 3 directors, one of each gender	2013
France	13.01.2011	Listed firms, private firms > 500 employees or > EUR 50m sales	BoD	20% each gender / 40% each gender; if > 8 directors, difference < 2	2014 / 2017
Italy	18.06.2011	Listed firms	BoD	20% each gender / 33.3% each gender	2012 / 2015
Belgium	14.09.2011	Listed firms	BoD	33.3% each gender	2017 (large), 2019 (small)
Malaysia	27.06.2011	Firms > 250 employees	BoD, SM	30% women	2016
India	30.08.2013 / 28.03.2018	Listed firms	BoD	1 woman / 1 independent woman	2015 / 2019 (large), 2020 (small)
Germany	06.03.2015	Listed firms subject to 50% co-determination law	SB	30% each gender	2016

Table 2: Descriptive statistics

This table reports descriptive statistics for the full sample of 2,462 firms (Panel A) and a sample comprising the 456 California-headquartered firms and 777 industry- and size-matched control firms (Panel B). The sample consists of one cross-section of Compustat firms as of the quota's adoption announcement (September 30). The sample includes all firms in Compustat with a data entry within one calendar year before September-end 2018, excluding utility and financial firms (SIC codes 4940-4949 and 6000-6999, respectively), firms with missing information on the state in which it is headquartered, firms headquartered outside the US, firms with negative book value of equity, firms with missing financial control variables, firms that only list American Depository Receipts, and firms without a listing on NYSE, AMEX, or NASDAQ. We also require at least 125 daily return observations during the 250-day estimation window that ends September 21 and complete return data for the entire five-day event window around the event date (October 1) and availability of board data from BoardEx. Panel A shows descriptive statistics for the resulting sample. Panel B shows descriptive statistics for a sample in which we draw, for each firm headquartered in California, the three closest size-matched non-California-headquartered firms in the same two-digit SIC code industry as control firms. While the same firm may serve as a matched control firm to more than one California-headquartered firm, every control firm is included only once in our sample. All financial ratios are winsorized at the 1st and 99th percentiles. Detailed variable definitions are in the Appendix.

Panel A: Full sample

Firm characteristic	Mean	P25	Median	P75	SD	N
CA HQ (d)	0.185	0.000	0.000	0.000	0.389	2,455
ROA	-0.019	-0.021	0.086	0.138	0.314	2,455
Leverage	0.221	0.010	0.201	0.365	0.196	2,455
Total assets (in millions)	6,350.820	135.192	709.906	2,941.623	28,228.046	2,455
PPE / TA	0.221	0.044	0.121	0.313	0.242	2,455
R&D / TA	0.084	0.000	0.008	0.092	0.160	2,455
Board size	8.110	7.000	8.000	9.000	2.153	2,455
Female directors (%)	0.153	0.000	0.143	0.222	0.126	2,455
2021 requ. failed (d)	0.840	1.000	1.000	1.000	0.367	2,455
# female directors missing	1.523	1.000	2.000	2.000	1.140	2,455
Shortfall (%)	0.214	0.100	0.222	0.333	0.164	2,455

Panel B: California-headquartered firms and industry- and size-matched control firms

Firm characteristic	Mean	P25	Median	P75	SD	N
CA HQ (d)	0.369	0.000	0.000	1.000	0.483	1,232
ROA	-0.087	-0.179	0.056	0.126	0.373	1,232
Leverage	0.180	0.000	0.130	0.314	0.185	1,232
Total assets (in millions)	4,432.676	80.387	381.679	1,454.982	22,241.947	1,232
PPE / TA	0.140	0.029	0.081	0.184	0.171	1,232
R&D / TA	0.131	0.000	0.055	0.171	0.196	1,232
Board size	7.748	6.000	8.000	9.000	1.980	1,232
Female directors (%)	0.146	0.000	0.143	0.222	0.124	1,232
2021 requ. failed (d)	0.875	1.000	1.000	1.000	0.331	1,232
# female directors missing	1.610	1.000	2.000	2.000	1.057	1,232
Shortfall (%)	0.231	0.111	0.250	0.333	0.156	1,232

Table 3: Market reaction to the quota announcement

This table reports differences in abnormal returns to the announcement of the gender quota in California between California-headquartered and non-California-headquartered control firms. Each column shows results from a pooled ordinary least squares regression of an abnormal return measure on a dummy variable set equal to one if a firm is headquartered in California (CA HQ (d)). Panel A reports the results from the base case model, in which we compute daily abnormal returns as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. For each California-headquartered sample firm, the three closest firms in terms of size that are active in the same two-digit SIC code industry are chosen as matched control firms. While the same firm may serve as a matched control firm to more than one California-headquartered firm in the treatment sample, every control firm is included only once in our sample. Across columns, we vary the length of the event window. In Panel B, we use a market-adjusted model, assuming that the beta of all stocks is equal to one. In Panels C and D, we use alternative matching procedures. In Panel C, we use propensity score matching to draw up to three matched control firms per California-headquartered sample firm. In Panel D, we use all non-California-headquartered firms as control firms. Panel E includes additional robustness tests for a two-day CAR that includes the event day and the subsequent day. In Columns 1 and 2, only the single closest industry- and size-matched firm or the closest propensity score matched control firm is selected into the control sample, respectively. In Column 3, we drop penny stocks with stock prices below USD 1. Columns 4 and 5 report results for a more or less restrictive winsorization of the dependent variable, respectively. Column 6 shows results from a median regression and no winsorizing of the dependent variable. The sample comprises all firms in Compustat with a data entry within one calendar year before September-end 2018, excluding utility and financial firms (SIC codes 4940-4949 and 6000-6999, respectively), firms with missing information on the state in which it is headquartered, firms headquartered outside the US, firms with negative book value of equity, firms with missing financial control variables, firms that only list American Depository Receipts, and firms without a listing on NYSE, AMEX, or NASDAQ. We also require at least 125 daily return observations during the 250-day estimation window and complete return data for the entire five-day event window around the event date (October 1) and availability of board data from BoardEx. If not otherwise noted, abnormal return measures are winsorized at the 1st and 99th percentiles. Standard errors are clustered at the two-digit SIC code industry level. All regressions include an intercept, which is not shown for brevity. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Panel A: Baseline results

Dependent variable:	AR(0)	AR(1)	CAR(0,1)	CAR(-1,1)	CAR(0,2)	CAR(-2,2)
	(1)	(2)	(3)	(4)	(5)	(6)
CA HQ (d)	-0.45***	-0.35**	-0.73***	-0.83**	-1.00***	-1.19***
	(-3.65)	(-2.15)	(-3.22)	(-2.30)	(-5.59)	(-2.99)
R^2	0.00	0.00	0.01	0.00	0.01	0.01
N	1,232	1,232	1,232	1,232	1,232	1,232

Panel B: Market-adjusted model

Dependent variable:	AR(0)	AR(1)	CAR(0,1)	CAR(-1,1)	CAR(0,2)	CAR(-2,2)
	(1)	(2)	(3)	(4)	(5)	(6)
CA HQ (d)	-0.41***	-0.35**	-0.72***	-0.75**	-1.01***	-1.02***
	(-3.75)	(-2.41)	(-3.39)	(-2.48)	(-5.05)	(-3.59)
R^2	0.00	0.00	0.01	0.00	0.01	0.01
N	1,232	1,232	1,232	1,232	1,232	1,232

Panel C: Three closest propensity score matched control firms

Dependent variable:	AR(0)	AR(1)	CAR(0,1)	CAR(-1,1)	CAR(0,2)	CAR(-2,2)
	(1)	(2)	(3)	(4)	(5)	(6)
CA HQ (d)	-0.32**	-0.36**	-0.64***	-0.88**	-0.77***	-1.11***
	(-2.52)	(-2.28)	(-2.93)	(-2.46)	(-3.43)	(-2.74)
R^2	0.00	0.00	0.00	0.01	0.00	0.01
N	1,208	1,208	1,208	1,208	1,208	1,208

Panel D: All non-CA headquarted firms as control firms

Dependent variable:	AR(0)	AR(1)	CAR(0,1)	CAR(-1,1)	CAR(0,2)	CAR(-2,2)
	(1)	(2)	(3)	(4)	(5)	(6)
CA HQ (d)	-0.57***	-0.47**	-0.99***	-1.12***	-1.08***	-1.25***
	(-3.62)	(-2.49)	(-3.44)	(-2.89)	(-3.76)	(-3.39)
R^2	0.01	0.00	0.01	0.01	0.01	0.01
N	2,455	2,455	2,455	2,455	2,455	2,455

Panel E: Additional robustness tests

Dependent variable:	CAR(0,1)					
Estimator:	OLS	OLS	OLS	OLS	OLS	Median regression
Control group:	Closest TA matches	Closest PS matches	Closest 3 TA matches			
Excluded:	-	-	Price <1$	-	-	-
Winsorizing:	1% & 99%	1% & 99%	1% & 99%	0.5% & 99.5%	5% & 95%	-
	(1)	(2)	(3)	(4)	(5)	(6)
CA HQ (d)	-0.52*	-0.51***	-0.60**	-0.91***	-0.62***	-0.63***
	(-1.78)	(-3.07)	(-2.46)	(-3.64)	(-3.23)	(-2.66)
R^2	0.00	0.00	0.00	0.01	0.01	
N	809	801	1,184	1,232	1,232	1,232

Table 4: Robustness tests: Control variables and alternative event dates

This table reports results from pooled ordinary least squares regressions of cumulative abnormal returns (CARs) on a dummy variable set equal to one if a firm is headquartered in California (CA HQ (d)), a set of financial control variables, and industry fixed effects. In Columns 1 to 4, CARs are estimated over two-day windows that include the event date and the first day after (i.e., October 1 and 2). In Columns 5 and 6, CARs are estimated over two-day windows that include the day of the introduction of the law (January 3) and the day after. In Columns 7 and 8, CARs are estimated over two-day windows that include the day of the succesfull Senate vote (May 31) and the day after. In Columns 9 and 10, CARs are estimate over three-day windows that include the day of the Assembly vote (August 29), the day of the second Senate vote (August 30), and the day after. Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The construction of the matched sample is described in detail in the caption of Table 2. Standard errors are clustered at the two-digit SIC code industry level in Columns 1, 2, and 5 to 10, at the one-digit SIC code level in Column 3, and at the three-digit SIC code level in Column 4. The CARs and all financial ratios are winsorized at the 1st and 99th percentiles. All regressions include an intercept, which is not shown for brevity. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Dependent variable:	CAR(0,1)				CAR(Jan. 3, Jan. 4)		CAR(May 31, Jun. 1)		CAR(Aug. 29, Aug. 31)	
Event:	Quota signed by Governor				Law introduced		Sucessfull Senate vote		Sucessfull Assembly vote and second Senate vote	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
CA HQ (d)	-0.62**	-0.61**	-0.64**	-0.57**	0.20	0.17	0.43	0.37	0.39	0.41
	(-2.41)	(-2.46)	(-2.79)	(-2.09)	(0.60)	(0.52)	(1.53)	(1.29)	(1.21)	(1.26)
ROA	0.32	0.40	0.31	0.40	1.72***	1.68***	1.48*	1.30	0.28	0.52
	(0.33)	(0.37)	(0.28)	(0.42)	(3.88)	(3.03)	(1.79)	(1.42)	(0.35)	(0.63)
Leverage	1.76**	1.42**	1.65*	1.12	0.96	0.82	0.37	0.80	-0.84	-0.82
	(2.45)	(2.06)	(2.25)	(1.33)	(1.60)	(1.46)	(0.29)	(0.59)	(-1.21)	(-1.16)
ln(Total assets)	0.05	0.03	0.07	0.04	-0.03	-0.01	-0.14*	-0.15*	-0.17*	-0.11
	(0.53)	(0.40)	(1.05)	(0.49)	(-0.47)	(-0.18)	(-1.69)	(-1.83)	(-1.99)	(-1.13)
PPE / TA	1.62	1.35	1.78	0.81	-0.43	-0.69	-0.33	-0.16	-0.27	-0.66
	(1.42)	(1.01)	(1.31)	(0.49)	(-0.58)	(-0.91)	(-0.54)	(-0.19)	(-0.17)	(-0.39)
R&D / TA	-0.14	0.38	0.24	0.72	4.62***	5.24***	2.50**	2.30**	1.70**	1.28
	(-0.15)	(0.51)	(0.54)	(0.65)	(5.85)	(4.95)	(2.57)	(2.08)	(2.19)	(1.58)
Industry FEs	-	2-SIC	1-SIC	3-SIC	-	2-SIC	-	2-SIC	-	2-SIC
R^2	0.02	0.09	0.04	0.14	0.02	0.05	0.01	0.04	0.01	0.03
N	1,232	1,232	1,232	1,232	1,184	1,184	1,220	1,220	1,229	1,229

Table 5: Difference-in-differences estimates using daily abnormal returns

This table reports results from pooled ordinary least squares regressions of daily abnormal returns (ARs) on a dummy variable set equal to one if a firm is headquartered in California (CA HQ (d)), a dummy set equal to one for observations measured after the implementation of the quota (Post (d)), and an interaction term between these two variables. The regression shown in Column 2 additionally includes industry fixed effects based on two-digit SIC codes, Column 3 additionally includes financial controls, and Column 4 includes firm fixed effects. The sample comprises four daily abnornmal stock return observations per firm, two before and two after the introduction of the quota. Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The construction of the matched sample is described in detail in the caption of Table 2. Standard errors are clustered at the firm level. Daily abnormal returns and all financial ratios are winsorized at the 1st and 99th percentiles. All regressions include an intercept, which is not shown for brevity. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Dependent variable:	AR(t)			
	(1)	(2)	(3)	(4)
CA HQ (d) × Post (d)	-0.38**	-0.38**	-0.38**	-0.38*
	(-2.14)	(-2.13)	(-2.13)	(-1.85)
Post (d)	-0.62***	-0.62***	-0.62***	-0.62***
	(-5.57)	(-5.55)	(-5.55)	(-4.82)
CA HQ (d)	-0.02	0.01	0.03	
	(-0.16)	(0.08)	(0.27)	
ROA			0.60*	
			(1.95)	
Leverage			0.07	
			(0.27)	
ln(Total assets)			0.05*	
			(1.75)	
PPE / TA			0.42	
			(0.90)	
R&D / TA			0.27	
			(0.47)	
Industry FEs	-	2-SIC	2-SIC	-
Firm FEs	No	No	No	Yes
R^2	0.02	0.03	0.04	0.30
N	4,928	4,928	4,928	4,928
Firms	1,232	1,232	1,232	1,232

Table 6: Robustness test: Accounting for the cross-sectional dependence of returns

This table reports cumulative abnormal stock returns for two portfolios, one comprising California-headquartered firms (CA HQ (d) = 1) and the other comprising a sample of industry- and size-matched non-California-head-quartered firms (CA HQ (d) = 0). The estimate for the daily abnormal return (AR) is obtained from estimating the following regression:

$$r_t = \alpha + ARd_t + \beta r_{wt} + \epsilon_t$$

where r_t is the daily equally-weighted portfolio return of all portfolio firms in excess of the 1-month U.S. treasury bill rate, d_t is a dummy variable set equal to one for observations in the event window and zero for observations in the estimation window, an r_{wt} is the daily value-weighted market index return in excess of the 1-month U.S. treasury bill rate. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The regression is estimated over a sample that includes all observations from the 250-day estimation window that ends on September 21 and the event window. Estimates for the cumulative abnormal returns are obtained by multiplying the obtained coefficient forr AR by the number of days in the event window. Differences in abnormal returns between California-headquartered and non-California-headquartered firms are obtained from estimating the regression above but with the dependent variable being the daily difference in portfolio returns of California-headquartered and non-California-headquartered firms. The construction of the matched sample is described in detail in the caption of Table 2. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

	CA HQ (d) = 1			CA HQ (d) = 0			Differences	
	Mean	t-value	Firms	Mean	t-value	Firms	Mean	t-value
AR (0)	-1.25%**	-2.53	455	-0.82%**	-2.05	777	-0.43%*	-1.73
AR (1)	-0.93%*	-1.89	455	-0.62%	-1.56	777	-0.31%	-1.26
CAR (0,1)	-2.18%***	-3.12	455	-1.44%**	-2.55	777	-0.74%**	-2.11
CAR (-1,1)	-2.16%**	-2.50	455	-1.39%**	-1.99	777	-0.77%*	-1.79
CAR (0,2)	-1.50%*	-1.72	455	-0.60%	-0.85	777	-0.90%**	-2.10
CAR (-2,2)	-1.66%	-1.47	455	-0.70%	-0.76	777	-0.96%*	-1.73

Table 7: Cross-sectional tests: Quota compliance and female board representation

This table reports results from pooled ordinary least squares regressions of two-day cumulative abnormal returns (CARs) on a dummy variable set equal to one if a firm is headquartered in California (CA HQ (d)), a variable related to quota compliance or female board representation, and an interaction term between these two. Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The construction of the matched sample is described in detail in the caption of Table 2. Standard errors are clustered at the two-digit SIC code level. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Dependent variable:	CAR(0,1)			
	(1)	(2)	(3)	(4)
CA HQ (d)	1.56**	0.15	0.25	-1.34***
	(2.07)	(0.30)	(0.53)	(-3.30)
CA HQ (d) × 2021 requ. failed (d)	-2.59***			
	(-3.91)			
2021 requ. failed (d)	0.52			
	(1.31)			
CA HQ (d) × # female directors missing		-0.54*		
		(-1.87)		
# female directors missing		0.06		
		(0.30)		
CA HQ (d) × Shortfall (%)			-4.18**	
			(-2.26)	
Shortfall (%)			0.84	
			(0.78)	
CA HQ (d) × Female directors (%)				4.14*
				(1.95)
Female directors (%)				-1.63
				(-1.52)
Constant	-2.18***	-1.82***	-1.92***	-1.49***
	(-5.43)	(-4.63)	(-5.47)	(-4.53)
R^2	0.01	0.01	0.01	0.01
N	1,232	1,232	1,232	1,232

Table 8: Cross-sectional tests: Board connections to potential female candidates

This table reports results from pooled ordinary least squares regressions of two-day cumulative abnormal returns (CARs) on a dummy variable equal to one if a firm is headquartered in California (CA HQ (d)), a proxy for the feminity of the networks of the directors of the focal firm, and an interaction term between these two. Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The construction of the matched sample is described in detail in the caption of Table 2. Standard errors are clustered at the two-digit SIC code level. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Dependent variable:	CAR(0,1)			
	(1)	(2)	(3)	(4)
CA HQ (d)	-3.55***	-2.78**	-3.01***	-2.82***
	(-3.07)	(-2.48)	(-2.74)	(-4.56)
CA HQ (d) × ln(# Female connections)	0.56**			
	(2.71)			
ln(# Female connections)	-0.32*			
	(-2.01)			
CA HQ (d) × ln(# Female employment connections)		0.46**		
		(2.11)		
ln(# Female employment connections)		-0.28***		
		(-3.33)		
CA HQ (d) × Female connections (%)			12.38*	
			(1.84)	
Female connections (%)			-8.63**	
			(-2.31)	
CA HQ (d) × Female employment connections (%)				11.90***
				(3.16)
Female employment connections (%)				-7.00**
				(-2.41)
Constant	-0.12	-0.49	-0.14	-0.50
	(-0.15)	(-1.23)	(-0.18)	(-0.74)
R^2	0.01	0.01	0.01	0.01
N	1,232	1,232	1,232	1,231

Table 9: Cross-sectional tests: Firm size and corporate governance

This table reports results from pooled ordinary least squares regressions of two-day cumulative abnormal returns (CARs) on a dummy variable set equal to one if a firm is headquartered in California (CA HQ (d)), a variable related to firm size or corporate governance, and an interaction term between these two. Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The construction of the matched sample is described in detail in the caption of Table 2. Standard errors are clustered at the two-digit SIC code level. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Dependent variable:	CAR(0,1)			
	(1)	(2)	(3)	(4)
CA HQ (d)	-2.22***	-0.89***	0.09	-0.76***
	(-3.93)	(-3.55)	(0.19)	(-4.04)
CA HQ (d) × ln(Total assets)	0.25***			
	(3.22)			
ln(Total assets)	0.07			
	(0.94)			
CA HQ (d) × Largest size quintile (d)		1.07**		
		(2.38)		
Largest size quintile (d)		1.68***		
		(4.60)		
CA HQ (d) × Modified E-Index			-0.37**	
			(-2.29)	
Modified E-Index			-0.09	
			(-0.92)	
CA HQ (d) × Sustainability index				1.59***
				(3.02)
Sustainability index				1.20**
				(2.65)
Constant	-2.15***	-1.92***	-1.85***	-1.77***
	(-3.86)	(-6.44)	(-4.71)	(-5.76)
R^2	0.01	0.03	0.01	0.02
N	1,232	1,232	1,091	769

Table 10: Spillover effects: The role of director supply

This table reports results from pooled ordinary least squares regressions of two-day cumulative abnormal returns (CARs) on measures of director supply for a sample of non-California-headquartered firms. Firms in this sample are selected by choosing the three closest firms in terms of size that are active in the same two-digit SIC code industry as the California-headquartered firms in our main sample. While the same firm may serve as a matched control firm to more than one California-headquartered firm, every control firm is included only once in the sample. Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The sample corresponds to the control sample used in Table 2 and subsequent tables. The construction of the sample thus follows the procedure used to construct the control sample, which is described in detail in the caption of Table 2. Standard errors are clustered at the two-digit SIC code level. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Dependent variable:	CAR(0,1)			
	(1)	(2)	(3)	(4)
ln(Missing female directors in industry in CA)	-0.24*			
	(-1.73)			
% missing female directors in industry in CA		-5.13***		
		(-2.99)		
Largest size quintile (d) × Female directors (%)			6.37**	
			(2.47)	
Female directors (%)			-3.43**	
			(-2.63)	
Largest size quintile (d)			0.54	
			(0.67)	
ln(Tenure female directors)				0.40**
				(2.56)
Constant	-0.76	-0.38	-1.46***	-2.42***
	(-1.17)	(-0.74)	(-4.24)	(-6.22)
R^2	0.01	0.02	0.02	0.00
N	777	774	777	548

Table 11: Spillover effects: The propensity to become subject to a future quota

This table reports results from pooled ordinary least squares regressions of two-day cumulative abnormal returns (CARs) on measures of political orientation of the state where a firm's headquarters are located (Columns 1 to 3) and measures that quantify the extent to which the state in which a firm is headquartered has followed California in legilslative issues in the past (Columns 4 to 7) for a sample of non-California-headquartered firms. Firms in this sample are selected by choosing the three closest firms in terms of size that are active in the same two-digit SIC code industry as the California-headquartered firms in our main sample. While the same firm may serve as a matched control firm to more than one California-headquartered firm, every control firm is included only once in the sample. Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The sample corresponds to the control sample used in Table 2 and subsequent tables. The construction of the sample thus follows the procedure used to construct the control sample, which is described in detail in the caption of Table 2. Standard errors are clustered at the two-digit SIC code level. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Dependent variable:	CAR(0,1)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
% votes Democrats	-3.92*						
	(-1.71)						
% votes Republican		4.94*					
		(1.98)					
Trump (d)			0.61*				
			(1.81)				
Soft female quota state (d)				-0.19			
				(-0.50)			
Medical cannabis legal (d)					-0.45		
					(-1.27)		
Cannabis legalization (score)						-0.30*	
						(-1.71)	
Quick medical cannabis follower (d)							-0.47*
							(-1.91)
Constant	0.25	-3.92***	-2.00***	-1.67***	-1.39***	-1.27***	-1.61***
	(0.19)	(-3.63)	(-6.93)	(-4.98)	(-3.11)	(-2.98)	(-5.09)
R^2	0.00	0.01	0.00	0.00	0.00	0.00	0.00
N	777	777	777	777	777	777	583

Table 12: Changes in board characteristics around the quota's adoption

This table reports results from pooled ordinary least squares regressions of different board characteristics on a dummy variable set equal to one if a firm is headquartered in California (CA HQ (d)), month dummy variables, and interaction terms between the California-headquarter dummy and the month dummy variables. For each California-headquartered firm and the sample of control firms, we compute the board characteristics for the end of September (the base month) as well as for the end of October, November, and December, respectively. In Column 1, we use the fraction of directors on the board that are female as dependent variable. In Column 2, we restrict the sample to firms that require at least one additional female director to fulfill the quota, and in Column 3, we restrict the sample to firms that require at least two additional female directors to fulfill the quota. In Columns 4 to 7, we use the sample from Column 1 but replace the dependent variable by a dummy equal to one if a firm at the end of a given month has no female director on the board (Column 4), by the fraction of female outside directors on the board (Column 5), by the fraction of female inside directors on the board (Column 6), and by the natural logarith of one plus board size multiplied by 100. The construction of the matched sample is described in detail in the caption of Table 2. All regressions include firm fixed effects and an intercept, which is not shown for brevity. Standard errors are clustered at the firm level. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Dependent variable:	Female directors (%)			No female (d)	Female outside directors (%)	Female inside directors (%)	ln(Board size)×100
Sample:	Full	# female directors missing > 0	# female directors missing > 1	Full	Full	Full	Full
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CA HQ (d) × October-end (d)	0.19	0.34*	0.49**	-0.02*	0.25	-0.11	0.61*
	(1.16)	(1.96)	(2.12)	(-1.94)	(1.36)	(-0.54)	(1.96)
CA HQ (d) × November-end (d)	0.32	0.45**	0.51*	-0.02**	0.37	-0.02	0.46
	(1.59)	(2.14)	(1.84)	(-2.02)	(1.62)	(-0.09)	(1.21)
CA HQ (d) × December-end (d)	0.45*	0.58**	0.69**	-0.03**	0.56**	-0.33	0.43
	(1.91)	(2.32)	(2.13)	(-2.53)	(2.14)	(-1.00)	(1.02)
October-end (d)	0.21***	0.21**	0.24**	-0.00	0.26***	0.11	0.01
	(2.65)	(2.34)	(2.09)	(-1.00)	(2.81)	(0.84)	(0.06)
November-end (d)	0.31***	0.34***	0.48***	-0.01	0.38***	0.02	0.22
	(2.91)	(3.12)	(3.40)	(-1.39)	(3.11)	(0.11)	(1.03)
December-end (d)	0.48***	0.53***	0.74***	-0.01*	0.53***	0.21	0.24
	(3.72)	(3.84)	(4.12)	(-1.70)	(3.62)	(0.81)	(1.02)
Firm FEs	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.02	0.03	0.04	0.01	0.02	0.00	0.00
N	4,903	4,289	2,858	4,903	4,903	4,798	4,903
Firms	1,232	1,075	715	1,232	1,232	1,232	1,232

Table 13: Director skills and experiences

This table reports results of tests for differences in means between the characteristics of different director subsets. Panel A reports descriptive statistics and results of comparison tests of newly appointed and incumbent directors at California-headquartered firms in the time period September 30 to December 31, 2018, for subsets of directors based on director gender. The sample of incumbent directors comprises all directors of the firms in the director appointment sample as of the quota announcement date. Panel B reports descriptive statistics and results of comparison tests of newly appointed and leaving directors at California-headquartered firms in the time period September 30 to December 31, 2018, for subsets of directors based on director gender. The bottom rows display sample sizes. Characteristics of incoming directors are determined as of the first day they are on the company's board, for leaving directors as of the last day of their board membership, and for incumbent directors as of the quota announcement date (September 30, 2018). The construction of the sample is described in detail in the caption of Table 2. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Panel A: Newly appointed female directors vs. incumbent female and male directors

	Newly appointed directors		Incumbent directors		Differences			
	Female (d)=1	Female (d)=0	Female (d)=1	Female (d)=0	(1) - (3)		(1) - (4)	
	(1)	(2)	(3)	(4)	(5)		(6)	
Age (yrs)	55.674	56.279	57.697	61.415	-2.023	-1.37	-5.741***	-3.43
Industry experience (d)	0.391	0.439	0.557	0.508	-0.166*	-1.80	-0.116	-1.51
Industry experience (yrs)	3.455	6.488	7.864	7.481	-4.408**	-2.42	-4.026**	-2.29
Director experience (d)	0.500	0.544	0.709	0.686	-0.209**	-2.37	-0.186**	-2.56
Director experience (yrs)	7.544	6.207	12.573	15.078	-5.029*	-1.82	-7.534**	-2.28
Outside director experience (d)	0.500	0.544	0.696	0.644	-0.196**	-2.21	-0.144*	-1.93
Outside director experience (yrs)	6.506	4.413	12.060	12.586	-5.554**	-2.12	-6.080**	-2.04
Current CEO (d)	0.000	0.043	0.013	0.065	-0.013	-0.76	-0.065*	-1.74
Current inside director (d)	0.023	0.043	0.013	0.084	0.009	0.38	-0.061	-1.43
Current C-level exec. (d)	0.136	0.152	0.067	0.111	0.070	1.27	0.026	0.50
Current VP-level exec. (d)	0.136	0.065	0.027	0.032	0.110**	2.34	0.104***	3.23
Outside director (d)	0.957	0.807	0.949	0.824	0.007	0.18	0.132**	2.32
Independent outside director (d)	0.705	0.587	0.960	0.932	-0.255***	-4.19	-0.228***	-5.09
Number of directors	46	57	79	449	125		495	
Number of outside directors	44	46	75	370	119		414	
Number of firms	42	43	49	77	69		77	

Panel B: Newly appointed female directors vs. appointed male directors and newly appointed female vs. leaving female directors

	Newly appointed directors		Leaving directors		Differences			
	Female (d)=1	Female (d)=0	Female (d)=1	Female (d)=0	(1) - (2)		(1) - (4)	
	(1)	(2)	(3)	(4)	(5)		(6)	
Age (yrs)	55.674	56.279	61.298	60.952	-0.605	-0.36	-5.624**	-2.54
Industry experience (d)	0.391	0.439	0.462	0.510	-0.047	-0.48	-0.070	-0.45
Industry experience (yrs)	3.455	6.488	4.254	7.510	-3.033*	-1.80	-0.799	-0.40
Director experience (d)	0.500	0.544	0.692	0.647	-0.044	-0.44	-0.192	-1.22
Director experience (yrs)	7.544	6.207	17.022	12.192	1.338	0.61	-9.478**	-2.11
Outside director experience (d)	0.500	0.544	0.692	0.598	-0.044	-0.44	-0.192	-1.22
Outside director experience (yrs)	6.506	4.413	14.296	10.133	2.094	1.19	-3.627**	-2.11
Current CEO (d)	0.000	0.043	0.000	0.048	-0.043	-1.40	0.000	0.00
Current inside director (d)	0.023	0.043	0.000	0.060	-0.021	-0.54	0.023	0.52
Current C-level exec. (d)	0.136	0.152	0.083	0.071	-0.016	-0.21	0.053	0.48
Current VP-level exec. (d)	0.136	0.065	0.000	0.000	0.071	1.12	0.136	1.35
Outside director (d)	0.957	0.807	0.923	0.824	0.150**	2.31	0.033	0.48
Independent outside director (d)	0.705	0.587	0.917	0.869	0.118	1.16	-0.212	-1.51
Number of directors	46	57	13	102	103		59	
Number of outside directors	44	46	12	84	90		56	
Number of firms	42	43	12	58	77		52	

Appendix: Variable definitions

This table reports variable definitions of all variables used in the paper as well as their data sources. Database mnemonics are in italics (if available).

Panel A: Firm characteristics

Variable	Definition	Source
CA HQ (d)	Dummy variable equal to one if a firm is headquartered (*state*) in California as of September-end 2018, zero otherwise.	Compustat
AR(t)	Abnormal return on day t, where October 1, 2018, marks the event date (t = 0). Daily abnormal returns are calculated as the observed return minus a predicted return. Except if otherwise noted, the predicted return is estimated using a market model regression where daily returns (adjusted for distributions and stock splits) are regressed on daily value-weighted index returns over a 250-day estimation window that ends six trading days prior to the event (September 21). At least 90 daily observations with non-missing stock and index return data are required. Winsorized at the 1% and 99% level.	Compustat
CAR(t_1,t_2)	Cumulative abnormal return, estimated as the sum of daily (unwinsorized) abnormal returns (AR) from t_1 to t_2 where October 1, 2018 marks the event date. Winsorized at the 1% and 99% level.	Compustat
ROA	Operating income before depreciation scaled by total assets (*oibdp* / *at*), winsorized at the 1% and 99% level.	Compustat
Leverage	Long-term debt and debt in current liabilities scaled by total assets ((*dltt* + *dlc*) / *at*), winsorized at the 1% and 99% level.	Compustat
Total assets	Total assets (*at*).	Compustat
PPE / TA	Property, plant, and equipment scaled by total assets (*ppent* / *at*), winsorized at the 1% and 99% level.	Compustat
R&D / TA	Research and development expenses scaled by total assets (*xrd* / *at*). Set to zero if *xrd* is missing in Compustat, winsorized at the 1% and 99% level.	Compustat
Board size	Number of directors on the board.	BoardEx
Female directors (%)	Fraction of directors on the board that are female.	BoardEx
Female inside directors (%)	Number of female inside directors to all inside directors on the board.	BoardEx
Female outside directors (%)	Number of female outside directors to all outside directors on the board.	BoardEx
2021 requ. failed (d)	Dummy variable equal to one if a firm fails to comply with the 2021 female director requirements of SB 826, that is, a firm does not have three female directors if board size is six or more, two female directors if board size is five, and one female director if board size is four or less, zero otherwise.	BoardEx
# female directors missing	Number of female directors necessary to fulfill the 2021 female director requirements of SB 826, that is, 3 minus the current number of female directors if board size is six or more, two minus the current number of female directors if board size is five, and 1 minus the current number of female directors if board size is four or less.	BoardEx
Shortfall (%)	# female directors missing scaled by board size.	BoardEx

No female (d)	Dummy variable equal to one if firm a has no female director on the board as of September-end 2018, zero otherwise.	BoardEx
# female connections	Number of distinct female directors that have a connection to at least one board member of the sample firm and sit on at least one board of another listed firm. A female director of another listed firm is defined to have a connection to a board member if she currently shares or shared in the past an overlapping work engagement at a firm other than the sample firm, graduated from the same university within one year, or is (was) active in the same social organization. For education connections, we disregard licenses (e.g., pilot license), leadership programs (e.g., Harvard's Advanced Management Program), certificates (e.g., CPA or CFA), and honorary degrees. For degrees, we require information on the degree year and that both graduate with a degree from the respective program. For social connections, we only consider "active roles", which we define as *rolename* not being a "member".	BoardEx
# female employment connections	Number of distinct female directors that have an employment connection to at least one board member of the sample firm and sit on at least one board of another listed firm. A female director of another listed firm is defined to have an employment connection to a board member if he or she currently shares or shared in the past an overlapping work engagement at a firm other than the sample firm.	BoardEx
Female connections (%)	Number of distinct female directors in the board's network scaled by the total number of directors in the board's network. A director of another listed firm is defined to have a connection to a board member if she currently shares or shared in the past an overlapping work engagement at a firm other than the sample firm, graduated from the same university within one year, or is (was) active in the same social organization. For education connections, we disregard licenses (e.g., pilot license), leadership programs (e.g., Harvard's Advanced Management Program), certificates (e.g., CPA or CFA), and honorary degrees. For degrees, we require information on the degree year and that both graduate with a degree from the respective program. For social connections, we only consider "active roles", which we define as *rolename* not being a "member".	BoardEx
Female employment connections (%)	Number of distinct female directors in the board's employment network scaled by the total number of directors in the board's employment network. A director of another listed firm is defined to have an employment connection to a board member if he or she currently shares or shared in the past an overlapping work engagement at a firm other than the sample firm.	BoardEx
Largest size quintile (d)	Dummy variable equal to one if a firm is in the highest Total asset quintile, zero otherwise.	Compustat
Modified E-Index	Score variable that adds one for each of the following five provisions in place as of September 30, 2018: Existance of a poison pill, existance of a staggered board, existance of a supermajority vote requirement to amend the bylaws, existance of a supermajority vote requirement to amend the charter, existance of a supermajority vote requirement for mergers. We disregard golden parachutes as the sixth provision of Bebchuk, Cohen, and Ferrell's (2009) E-Index because it is not available in SharkRepellent.	Shark-Repellent

Sustainability index	For each of the five KLD categories “Community”, “Employee Relations”, “Environment”, “Human Rights”, and “Product”, we compute a positive and negative index based on the positive (“Strengths”), or negative (“Concerns”) indicator variables. As the number of strengths and concerns varies across categories, we scale the strengths and concerns for each category to obtain two indices that range from 0 to 1. We then divide the number of strengths (concerns) for each firm-year within each category by the maximum possible number of strengths (concerns) in each category-year. Within each category, in each firm-year, we subtract the concerns index from the strengths index to end up with a net sustainability score that ranges from –1 to +1. We sum up the five different category indices, which yields our sustainability score that ranges from -5 to +5 with a higher score indicating a more sustainable firms. We use the mean sustainability score of all available observations on each sample firm, a maximum of seven yearly observations (2010 – 2016).	KLD Stats
Missing female directors in industry in CA	Sum of # female directors missing across all California-headquartered firms in two-digit SIC code industry.	Compustat/ BoardEx
% missing female directors in industry in CA	Sum of # female directors missing across all California-headquartered firms in two-digit SIC code industry scaled by the sum of # female directors missing across all firms in two-digit SIC code industry.	Compustat/ BoardEx
Tenure female directors	Mean tenure in years of all female directors on the board. Set to missing if firm has no female director.	BoardEx
% votes Democrats	Fraction of votes obtained by Democratic Party in 2016 Presidential Election in the state where a company’s headquarter is located.	Politico
% votes Republican	Fraction of votes obtained by Republican Party in 2016 Presidential Election in the state where a company’s headquarter is located.	Politico
Trump (d)	Dummy variable equal to one if the state where a company’s headquarter is located was won by the Republican Party in the 2016 Presidential Election, zero otherwise.	Politico
Soft quota state (d)	Dummy variable equal to one if the state where a company’s headquarter is located has introduced non-binding board gender legislation (Illinois (May 2015), Massachusetts (October 2015), Ohio (April 2016), Colorado (March 2017), and Pennsylvania (April 2017)), zero otherwise.	State-level legislative websites
Medical cannabis legal (d)	Dummy variable equal to one if the firm is headquartered in a state that has legalized cannabis for medical use, and zero otherwise. States that have only legalized the use of Cannabidiol (CBD) oils are coded as zero because some of these states banned the active chemical ingredient, Tetrahydrocannabinol (THC) (e.g., Kansas) or set THC limits close to zero (e.g., Indiana), resulting in legal uncertainty amongst sellers and significantly reducing the availability of CBD oils as they can contain traces of THC.	State-level legislative websites
Cannabis legalization (score)	Count variable that is equal to zero if any type of cannabis use is considered illegal, equal to one if the use of cannabis is legal for medical purposes (again excluding the exclusive use of CBD oil), equal to two if the recreational consumption of cannabis is illegal, but has been decriminalized, and equal to three if the recreational consumption of cannabis is legal.	State-level legislative websites

Quick medical cannabis follower (d)	Dummy variable equal to one if the state where a company's headquarter is located has legalized the use of cannabis for medical treatment below the median time of legalization following California, zero otherwise.	State-level legislative websites

Panel B: Director characteristics

Variable	Definition	Source
Female (d)	Dummy variable equal to one if a director's gender (*gender*) is female, zero if male.	BoardEx
Age (yrs)	Years passed since the date of birth (*dob*). If only month and year of *dob* is available, we use the 15th of a given month, if only the year of *dob* is available, we use July 1st of a given year.	BoardEx
Current CEO (d)	Dummy variable equal to one if an outside director serves as a CEO of another listed firm, and zero otherwise.	BoardEx
Current inside director (d)	Dummy variable equal to one if an outside serves as an inside director of another listed firm, and zero otherwise.	BoardEx
Current C-level exec. (d)	Dummy variable equal to one if an outside serves as a C-level executive of another listed firm, and zero otherwise. C-level executives include titles such as Chief Executive Officer, Chief Financial Officer, Chief Operating Officer, etc.	BoardEx
Current VP-level exec. (d)	Dummy variable equal to one if an outside director serves as a Vice President of another listed firm, and zero otherwise. VP executives include Vice Presidents and Senior Vice Presidents.	BoardEx
Industry experience (d)	Dummy variable equal to one if a director has experience at another listed firm that is active in the same two-digit SIC code industry as the sample firm and zero otherwise.	BoardEx/ CRSP
Industry experience (yrs)	Sum of the duration in years across all positions at other listed firms that are active in the same two-digit SIC code industry as the sample firm.	BoardEx/ CRSP
Director experience (d)	Dummy variable equal to one if a director has experience as a director of another listed firm and zero otherwise.	BoardEx
Director experience (yrs)	Sum of the duration in years across all director positions at other listed firms.	BoardEx
Outside director experience (d)	Dummy variable equal to one if a director has experience as an outside director of another listed firm and zero otherwise.	BoardEx
Outside director experience (yrs)	Sum of the duration in years across all outside director positions at other listed firms.	BoardEx
Outside director (d)	Dummy variable equal to one if a director is an outside director (*ned*), zero if not.	BoardEx
Independent outside director (d)	Dummy variable equal to one if an outside director is independent (*rolename*), zero if not.	BoardEx

Table OA.1: Balancing tests

This table reports differences in firm characteristics between California-headquartered firms (CA HQ (d) = 1) and non-California-headquartered U.S. firms (CA HQ (d) = 0). The table reports means, medians, and the number of observations. The table also reports results from tests for differences in means and medians in firm characteristics between the two subsamples. The sample comprises all firms in Compustat with a data entry within one calendar year before September-end 2018, excluding utility and financial firms (SIC codes 4940-4949 and 6000-6999, respectively), firms with missing information on the state in which it is headquartered, firms headquartered outside the US, firms with negative book value of equity, firms with missing financial control variables, firms that only list American Depository Receipts, and firms without a listing on NYSE, AMEX, or NASDAQ. We also require at least 125 daily return observations during the 250-day estimation window that ends September 21 and complete return data for the entire five-day event window around the event date (October 1) and availability of board data from BoardEx. Panel A shows balancing tests for the resulting sample. Panel B shows balancing tests for a sample in which we draw, for each firm headquartered in California, the three closest firms in terms of size that are active in the same two-digit SIC code industry. While the same firm may serve as a matched control firm to more than one California-headquartered firm, every control firm is included only once in our sample. All financial ratios are winsorized at the 1st and 99th percentiles. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Panel A: Control firms are all non-CA headquartered firms

	CA HQ (d) = 1			CA HQ (d) = 0			Differences			
	Mean	Median	N	Mean	Median	N	Mean	t-value	Median	z-value
ROA	-0.129	0.016	455	0.006	0.092	2,000	-0.135***	-8.41	-0.076***	-8.79
Leverage	0.168	0.114	455	0.233	0.219	2,000	-0.065***	-6.45	-0.105***	-6.63
Total assets	5,495.984	384.506	455	6,545.295	859.861	2,000	-1,049.311	-0.72	-475.355***	-6.31
PPE / TA	0.121	0.067	455	0.244	0.144	2,000	-0.123***	-9.97	-0.077***	-10.44
R&D / TA	0.166	0.100	455	0.065	0.002	2,000	0.101***	12.55	0.099***	15.54
Board size	7.664	8.000	455	8.211	8.000	2,000	-0.547***	-4.92	-0.000***	-4.81
Female directors (%)	0.146	0.143	455	0.155	0.143	2,000	-0.009	-1.30	-0.000	-1.24
2021 requ. failed (d)	0.888	1.000	455	0.829	1.000	2,000	0.059***	3.12	0.000***	3.12
# missing female directors	1.644	2.000	455	1.496	2.000	2,000	0.148**	2.51	0.000**	2.08
Shortfall (%)	0.236	0.250	455	0.210	0.222	2,000	0.026***	3.04	0.028***	3.02

Panel B: Control firms from same two-digit SIC industry, closest 3 total assets matches

	CA HQ (d) = 1			CA HQ (d) = 0			Differences			
	Mean	Median	N	Mean	Median	N	Mean	t-value	Median	z-value
ROA	-0.132	0.016	455	-0.060	0.069	777	-0.072***	-3.27	-0.053***	-3.96
Leverage	0.168	0.114	455	0.187	0.140	777	-0.019*	-1.74	-0.025*	-1.85
Total assets	5,495.984	384.506	455	3,810.019	376.335	777	1,685.965	1.28	8.171	0.18
PPE / TA	0.120	0.067	455	0.151	0.090	777	-0.031***	-3.06	-0.023***	-3.31
R&D / TA	0.169	0.100	455	0.109	0.027	777	0.06***	5.28	0.073***	7.59
Board size	7.664	8.000	455	7.798	8.000	777	-0.134	-1.15	-0.000	-1.10
Female directors (%)	0.146	0.143	455	0.146	0.143	777	-0.000	-0.00	-0.000	-0.03
2021 requ. failed (d)	0.888	1.000	455	0.867	1.000	777	0.02	1.05	0.000	1.05
# missing female directors	1.644	2.000	455	1.591	2.000	777	0.053	0.85	0.000	0.66
Shortfall (%)	0.236	0.250	455	0.228	0.250	777	0.007	0.78	0.000	0.66

Table OA.2: Univariate return differences

This table reports differences in abnormal returns between California-headquartered firms and U.S. non-California-headquartered firms around the quota adoption announcement. We report results from t-tests against zero for different abnormal return measures for the subsample of firms headquartered in California (CA HQ (d) = 1) and for the subsample of firms headquartered in any other U.S. state but California (CA HQ (d) = 0). Daily abnormal returns are computed as the observed return minus the predicted return from a market model regression estimated over a 250-day estimation window that ends on Friday, September 21. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. The sample comprises all firms in Compustat with a data entry within one calendar year before September-end 2018, excluding utility and financial firms (SIC codes 4940-4949 and 6000-6999, respectively), firms with missing information on the state in which it is headquartered, firms headquartered outside the US, firms with negative book value of equity, firms with missing financial control variables, firms that only list American Depository Receipts, and firms without a listing on NYSE, AMEX, or NASDAQ. We also require at least 125 daily return observations during the 250-day estimation window and complete return data for the entire five-day event window around the event date (October 1) and availability of board data from BoardEx. Panel A reports results from comparing abnormal return measures of California-headquartered firms to a sample of matched non-California-headquartered firms. The matched sample is obtained by choosing for each California-headquartered sample firm the three closest firms in terms of size that are active in the same two-digit SIC code industry. While the same firm may serve as a matched control firm to more than one California-headquartered firm, every control firm is included only once in our sample. Panel B reports results from comparing abnormal return measures of California-headquartered firms to all U.S. non-California-headquartered firms. All abnormal return measures are winsorized at the 1st and 99th percentiles. Detailed variable definitions are in the Appendix. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively.

Panel A: Control firms from same two-digit SIC industry, closest 3 total assets matches

	CA HQ (d) = 1			CA HQ (d) = 0			Differences	
	Mean	t-value	N	Mean	t-value	N	Mean	t-value
AR (0)	-1.43%***	-9.67	455	-0.98%***	-7.87	777	-0.45%**	-2.29
AR (1)	-1.06%***	-7.60	455	-0.72%***	-6.29	777	-0.35%*	-1.90
CAR (0,1)	-2.46%***	-11.56	455	-1.73%***	-10.10	777	-0.73%***	-2.65
CAR (-1,1)	-2.64%***	-9.96	455	-1.80%***	-8.56	777	-0.83%**	-2.44
CAR (0,2)	-1.85%***	-7.35	455	-0.85%***	-3.88	777	-1.00%***	-2.92
CAR (-2,2)	-2.39%***	-7.00	455	-1.20%***	-4.49	777	-1.19%***	-2.73

Panel B: Control firms are all non-CA headquartered firms

	CA HQ (d) = 1			CA HQ (d) = 0			Differences	
	Mean	t-value	N	Mean	t-value	N	Mean	t-value
AR (0)	-1.44%***	-9.94	455	-0.87%***	-13.74	2,000	-0.57%***	-3.77
AR (1)	-1.09%***	-8.04	455	-0.62%***	-10.54	2,000	-0.47%***	-3.35
CAR (0,1)	-2.46%***	-11.70	455	-1.48%***	-16.16	2,000	-0.99%***	-4.55
CAR (-1,1)	-2.61%***	-10.29	455	-1.49%***	-14.03	2,000	-1.12%***	-4.40
CAR (0,2)	-1.86%***	-7.47	455	-0.77%***	-6.95	2,000	-1.08%***	-4.14
CAR (-2,2)	-2.31%***	-7.23	455	-1.06%***	-7.84	2,000	-1.25%***	-3.89