As California goes, so goes the nation? Board gender quotas and the legislation of non-economic values*

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Abstract

In 2018, California became the first U.S. state to introduce a mandatory board gender quota for all firms headquartered in the state. We document negative announcement returns to the adoption of the quota for Californian firms, but also large negative spillover effects on a matched group of non-Californian firms, particularly those located in states that followed California's legislative lead in the past by raising minimum wages or legalizing cannabis. Frictions on the director labor market only explain a small fraction of value losses of Californian firms. They do not explain the negative spillover effects on firms in other states. We propose shareholders' fear of further legislation of non-economic values as a new explanation for the negative announcement returns to gender quotas. In line with this view, we find that firms with higher policy sensitivity show the strongest reaction.

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1 Introduction

Women are still heavily underrepresented in corporate leadership positions. Numbers based on the Russell 3000 index, which represents approximately 98% of U.S. equity market capitalization, show that as of June 2018, women held 16.9% of director positions (see Figure 1). Furthermore, the fraction of female directors in the U.S. is growing slowly, at an average annual rate of 0.8%. If no proactive measures are taken and the current rate of growth remains unchanged, it would take more than 40 years to achieve gender parity at U.S. boards.[1]

Several countries have responded to gender inequality in boardrooms by adopting mandatory quotas.[2] The first country to act was Norway, which introduced a gender quota of 40% female representation in 2003. The adoption of the quota was associated with large value losses for affected firms (Ahern and Dittmar (2012)). Following Norway's lead, Belgium, France, Germany, Iceland, India, Israel, Italy, and Portugal have all established similar quotas (see Table 1).

In the U.S., California is the first state to adopt a mandatory board gender quota. On September 30, 2018, Governor Brown signed Senate Bill 826 into law, which requires that all national and foreign companies headquartered in California with a listing at a major U.S. stock exchange 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 with comparably weak penalties, including a payment of $100,000 for the first violation and $300,000 for each subsequent violation. Furthermore, it is still an open question whether the quota is legal, as it may conflict with the corporate internal affairs doctrine as well as California and U.S. federal civil rights laws (Grundfest (2018)).

Therefore, one may not expect large reactions to the adoption of the quota given that it should be quite easy for Californian firms to buy themselves out by paying a rather low penalty or by betting on the quota being confirmed as unconstitutional. However, in this paper we not only document a robust and significantly negative stock market reaction of Californian firms to the

[1]For an overview of potential reasons for the low fraction of women in leadership, see Bertrand (2017).
[2]Adams (2016) provides an overview on worldwide board gender diversity policies.

introduction of the gender quota, but also find large negative spillover effects on non-Californian firms. While announcement returns of Californian firms are -2.25% on average for a two-day event window, non-Californian firms experience a negative announcement return of -1.53%. This is remarkable given that the latter are not directly affected by the introduction of the gender quota in California. In economic terms, the mean reduction in shareholder value amounts to $84.73 million in non-Californian and $285.94 million in Californian firms, respectively (based on the mean market capitalization in the two groups). This value loss even increases in economic magnitude for larger event windows.

Why is there such a large value loss for both Californian and non-Californian firms? The predominant explanation put forward in prior literature is that there may be frictions in the director labor market that make it costly for firms to hire additional female directors (Ahern and Dittmar (2012)). For example, there may not be enough qualified female candidates and/or the pool of female directors may be costly to access if female directors are part of different job networks to which firms don't have regular access (Calvó-Armengol and Jackson (2004); Beaman, Keleher, and Magruder (2018)). Consistently, Ferreira, Ginglinger, Laguna, and Skalli (2017) show that French firms changed their director search technology in response to the introduction of the French board gender quota in 2011. In this case, firms have to establish costly new recruiting procedures to identify suitable female candidates and may have to pay higher wages to these scarce candidates.[3]

Frictions in the director labor market should, however, affect Californian firms more strongly as these firms are directly affected by the introduction of the gender quota. They should be less relevant for non-Californian firms: Although they may expect to face similar regulation in the future, they are not yet directly affected by a gender quota. In line with this view, we find that Californian firms facing a larger friction, i.e., firms with a larger female director gap on their board, indeed react more strongly to the gender quota announcement. However, we find that even after controlling for frictions in the director labor market, there remains a large and

[3]An alternative explanation would be that shareholders expect too much monitoring if female directors are added to the board and, as a result, a reduction of firm value (Adams and Ferreira (2007; 2009)).

unexplained negative announcement return for Californian firms. In addition, frictions do not explain the negative spillover effects on non-Californian firms.

Therefore, we propose a new channel by which the negative valuation effects of Californian and non-Californian firms in response to the gender quota can be explained: Shareholders' expectations that further non-economic values will be legislated and imposed on firms. For example, shareholders may interpret a gender quota as a signal that further (costly) legislation regarding equal opportunities (or targeting other ESG-related aspects) may follow. If shareholders think that such laws will not increase cash flows beyond the additional costs from implementing the new laws, stock prices will decline.

In recent years, there have been several attempts to legislate non-economic values. At the federal level, for instance, Senator Elizabeth Warren proposed the Accountable Capitalism Act, a federal bill that would require that employees elect 40% of a board of directors of any corporation with over $1 billion in tax receipts. At the state level, California has frequently been the first state to enact non-economic value legislation, which was often later adopted by other U.S. states. An example of California's leadership in legislating non-economic values is the adoption of a statewide $12 minimum hourly wage (with staged increases to $15 by January 1, 2022), which will more than double the federal minimum wage of $7.25.

There are several papers showing that the uncertainty associated with possible changes in government policy has implications for firm values (e.g., Pástor and Veronesi (2012; 2013); Brogaard and Detzel (2015); Kelly, Pástor, and Veronesi (2016); Koijen, Philipson, and Uhlig (2016)). If the introduction of a gender quota affects shareholders' general expectation that further non-economic values will be imposed on firms through government legislation, negative valuation effects may arise for both Californian and non-Californian firms.

To test this conjecture, we use the Economic Policy Uncertainty (EPU) index of Baker, Bloom, and Davis (2016) to compute a firm-level measure of sensitivity to the firm's regulatory environment. Controlling for the impact of frictions, we find that Californian firms that are more sensitive to economic policy uncertainty indeed experience more negative announcement returns than other Californian firms. Non-Californian firms in democratic states and in states that followed California's legislation in the past, for instance, by introducing minimum wages exceeding the federal level or by legalizing cannabis, also show more negative announcement returns. Mirroring the California result, we find that non-Californian firms that are sensitive to economic policy uncertainty react stronger in states more likely to follow California's legislative lead. We conclude from our results that the large value losses of Californian and non-Californian firms are unlikely to be explained solely by the direct costs of a gender quota, resulting from frictions on the director labor market. In addition, shareholders seem to expect a higher likelihood of future non-economic value legislation. Expected costs associated with this type of legislation seem to be priced at the time when a law targeting a non-economic value, such as board diversity, is introduced.

Our results provide a new explanation for the large and negative announcement effects that have been documented for the introduction of gender quotas (e.g., Ahern and Dittmar (2012); Matsa and Miller (2013), Greene, Intintoli, and Kahle (2019); Hwang, Shivdasani, and Simintzi (2019)). These studies assign the negative quota effect to frictions associated with the appointment of female directors to the board. We show that frictions in the director labor market only partially explain the overall announcement effect and propose that the effect is partly explained by a more general fear that the quota may only be the beginning of legislation of this type. Our findings imply that non-economic laws with respect to the corporation have economic effects. Shareholders seem to assess legislation on gender equality as portending to future legislation which will affect (and reduce) the economic value of the firm, at least with respect to shareholder wealth. This is particularly relevant in times where the purpose of the corporation is under reconsideration in the U.S., with CEOs of major U.S. companies announcing in August 2019 that they are committed to lead their companies for the benefit of all stakeholders, and not just maximization of shareholder value.[4]

[4]The full statement is available at https://www.businessroundtable.org/business-roundtable-redefines-the-purpose-of-a-corporation-to-promote-an-economy-that-serves-all-americans. See also, for example, David Gelles and David Yaffe-Bellany, *Shareholder Value Is No Longer Everything, Top C.E.O.s Say*, The New York Times, Aug. 19, 2019.

2 The gender quota in California: Senate Bill 826

In September 2013, California Senate Concurrent Resolution 62 was passed by both houses of the California state legislature. The resolution was non-binding, and required that: "within (...) three years, every publicly held corporation in California with nine or more director seats have a minimum of three women on its board, every publicly held corporation in California with five to eight director seats have a minimum of two women on its board, and every publicly held corporation in California with fewer than five director seats have a minimum of one woman on its board". The three-year period specified in the resolution ended on December 31st, 2016. According to our own estimations, approximately 194 (44%) of the Russell 3000 firms headquartered in California failed to comply with the resolution's target number of female directors by that date. Among the 446 publicly traded California-headquartered firms included in the Russell 3000 index, female directors held 526 (14%) seats, men held 3,090 (86%) seats, and 126 (28%) firms had no female director.

Senate Bill 826 ("SB 826") was introduced on January 3rd, 2018, with a stated purpose of addressing the continued deficit in female directors at publicly traded firms in California. It first passed the Senate on May 31st, 2018, (22:6:11 votes) and the Assembly on August 29th, 2018 (41:13:26 votes). The Senate passed the amendments made by the Assembly on August 30th, 2018 (23:8:9 votes).

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. The Bill requires that these firms have a minimum of one female director on its board by the end of 2019. The Bill also prescribes that, by the end of calendar year 2021, Californian companies with five directors must have two female directors, and companies with six or more directors must have three female directors.

The Bill provides an enforcement mechanism which subjects non-compliant companies to financial penalties. For a first violation, a fine of $100,000 is imposed; for a second or subsequent violation, the fine increases 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.

The Bill was presented to the Governor's office for signing on September 10th, 2018. Governor Brown is known for his willingness to veto bills (he vetoed 12% of bills in 2017 and 15% in 2016).[5] He did not initially announce what he would do with Senate Bill 826, and contemporary news and commentary indicated that it was uncertain whether he would sign the Bill.[6] On Sunday September 30th, 2018, he signed SB 826 into law and announced it on the same day.

Governor Brown signed the Bill despite significant commentary that it could be unconstitutional (Grundfest (2018)). The argument for unconstitutionality has thus far focused on two grounds. First, the Bill may violate the internal affairs doctrine since it purported to regulate firms incorporated in Delaware or another State but headquartered in California. Second, the imposed quota may violate equal rights protections of the U.S. and Californian constitutions. As of the date of this paper, two lawsuits have been filed (one in federal and one in state court) challenging the law on the grounds that it violates the equal protection clause of the U.S. constitution.[7]

While several European countries have introduced gender quotas (see Table 1), California is the first state in the U.S. to adopt such a binding resolution. Legislators in five states (Illinois, Massachusetts, New Jersey, New York, and Washington) proposed legally binding board gender quotas similar to SB 826 – however, none of them have passed at the time of this writing.

[5]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 New York Times, Oct. 1, 2018 (Governor Brown has a "willingness to wield the veto pen").

[6]See also 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 . . . .").

[7]See Andrew Sheeler, *California man sues to overturn 'woman quota' in state gender equity law*, Sacramento Bee, Nov. 13, 2019.

3 Data and sample selection

3.1 Sample selection and data sources

To construct our sample, we select all firms in Compustat with a data entry within one calendar year before the date of the introduction of the gender quota in California (end of September, 2018). We drop utility and financial firms, firms with missing information on the state in which they are headquartered, firms headquartered outside the U.S., firms with negative book value of equity, and firms with missing financial control variables. As the quota law only applies to firms headquartered in California "with outstanding shares listed on a major United States stock exchange", we additionally drop firms that only list American Depository Receipts and firms without a listing on NYSE, AMEX, or NASDAQ. We supplement this sample with stock price data from Compustat.[8] If Compustat reports more than one stock price series for a firm, we choose the time-series with the highest market capitalization as of the event date among those time-series with sufficient data to estimate the market model. We also require data on the board of directors of our sample firms, which we obtain from BoardEx. Our final sample consists of 2,455 firms, out of which 455 (18.5%) are headquartered in California and 2,000 are headquartered in other states of the U.S. or the D.C. We describe all variables used in our empirical analyses in detail in the Appendix of the paper.

Governor Brown signed the law on Sunday, September 30th, 2018, and the adoption of the law was publicly announced on the same day. We define the event date as the first trading day after the announcement: Monday, October 1st. The choice of this event date is justified by the patterns reported in Figure 2, which displays the distribution of newspaper coverage on the gender quota in California. The weekly distribution of articles on the gender quota show that newspaper coverage is concentrated in the week following the signature of the bill by Governor Brown on Sunday, September 30. Reading a random subset of articles published before

[8] Note that at the time of writing, daily stock return data from the Center for Research in Security Prices (CRSP) and the CRSP-Compustat matching table were not yet available for our sample period. 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. Only 326 out of 50,471 return observations (0.6%) differ by more than 0.01 percentage points, suggesting that return data in CRSP and Compustat are very similar.

September 30 confirms that there was considerable uncertainty whether the Governor would sign the controversial bill.

We compute daily abnormal returns (ARs) as the observed return less the predicted return from a market model regression[9] estimated over a 250-day estimation window that ends on Friday, September 21st, i.e., six trading days before the event date. 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 a symmetric five-day window around the event date. 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 reaction to the introduction of the gender quota is based on a two-day event window, which includes the event day (October 1st) and the following day (October 2nd). All abnormal returns are winsorized at the 1st and 99th percentiles to mitigate the effect of outliers.

3.2 Descriptive statistics

Panel A of Table 2 reports descriptive statistics for the 455 firms subject to California's board gender quota. As of September 30th, these firms have a mean (median) board size of 7.7 (8.0), out of which 14.6% (14.3%) are female directors. The quota imposed on firms headquartered in California mandates that firms must have, by the end of calendar year 2021, three female directors if a board comprises six or more directors, two female directors if a board comprises five directors, and one female director if there are four or less directors on a board. Using data on board size and directors' gender, we find that, at the time of the adoption of the quota law, 88.8% of all firms do not comply with the mandated quota that would apply to their current board size. To comply with the quota, Californian firms would have to appoint on average 1.7 (median: 2.0) female directors (# missing female directors).

[9] Our results are very similar if we use a market-adjusted model, which assumes a beta of one for every firm (see Section 2.1 and Table OA.1 in the Online Appendix).

3.3 Construction of matched control sample

To obtain treatment and control groups that are similar in terms of observable and thus presumably unobservable firm characteristics, we choose, for each of our 455 treatment firms headquartered in California, three non-California-headquartered firms that share the same primary two-digit SIC code and are closest in terms of total assets. We match with replacement, i.e., a firm in the control sample may serve as such for more than one treatment firm, but we include every control firm only once in the sample. The resulting matched sample comprises 1,232 firms, 455 (36.9%) in the treatment group and 777 in the control group. Descriptive statistics for the matched control firms are displayed in Panel B of Table 2 and are very similar to those of the Californian firms reported in Panel A, i.e. they are not significantly different from each other. Note that the control sample is not dominated by one state or a small number of states: The largest fraction of control firms is headquartered in Massachusetts (14.03%), followed by Texas (10.68%) and New York (10.68%). In total, the control group includes firms that are headquartered in 42 different states and the D.C.

4 The impact of the gender quota on firm performance

To analyze firms' stock price reaction to the introduction of the quota, we regress different abnormal return measures on a dummy variable equal to one if a firm's headquarters are located in the state of California, and zero otherwise. The event date is Monday, October 1st, 2018, the first trading day after the public announcement of the adoption of the quota by the Governor's office.

The identifying assumption central to a causal interpretation of such a difference-in-differences analysis is that treated and control firms share parallel trends before the onset of treatment, i.e., the introduction of the quota on September 30th 2018. In Table 3, we provide evidence in support of the parallel trends assumption: The results show that for four of five market-adjusted run-up returns, the California and non-California firms do not differ in the 50 to 250 days leading up to the quota adoption. Only the difference in 200-day run-up returns is (borderline) significant, but here – in contrast to our event study results – Californian firms outperform non-Californian firms by a modest 2.42% over 200 trading days.

Table 4 reports results from regressions of six different abnormal return measures, estimated over different event windows that range from one to five days in length, on the California-headquarters dummy variable. Results in column (1) show that the abnormal return on the first trading day after the announcement is 0.45% lower for Californian firms than for the matched non-Californian control firms. The coefficient on the California dummy variable is significant at the 5% level. Results in column (2) indicate that, also on the second trading day after the announcement, Californian firms significantly underperform the sample of matched control firms by -0.32%. When we compute cumulative abnormal returns for the first two event days (column (3)), returns are 0.72% lower for Californian firms in comparison to the sample of matched control firms.[10] In our further analysis, we use the regression model in column (3) as our baseline specification.

In economic terms, this estimate translates into a two-day average loss in shareholder value of around $91.38 million at Californian versus non-Californian firms. Results remain statistically significant and coefficient are even larger 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)).

The constant is negative and statistically significant across all specifications. In our baseline specification in column (3), we find that non-Californian firms experience a negative announcement return of -1.53% on average upon the adoption of the quota in California. Accounting for the relative underperformance of Californian versus non-Californian firms of -0.72%, the negative announcement return of Californian firms is thus -2.25% on average. Based on the mean market capitalization, the mean reduction in shareholder value amounts to $84.73 million for non-Californian and $285.94 million for Californian firms, respectively. Hence, while Californian firms react more strongly to the adoption of the board gender quota in California, non-Californian

[10] We obtain economically identical results if we run a classical difference-in-differences regression using daily abnormal returns as dependent variable and a post dummy, a California dummy, and an interaction term between the two. For the results, see Section 2.2 and Table OA.2 in the Online Appendix.

firms, which are not directly affected by the quota, show negative and significant announcement returns as well. We explore various explanations for the negative announcement returns at Californian and non-Californian firms in the next section.

In robustness tests, we analyze whether our estimate for the loss in shareholder value around the law's announcement is sensitive to controlling for firm-level covariates. 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 4. In column (1) of Table OA.4 in the online appendix, we re-estimate the baseline regression reported in column (3) of Table 4 and augment it with a set of firm-specific control variables. The results in column (1) of Table OA.4 show that the coefficient on the dummy indicating whether a firm is headquartered in California remains economically similar to Table 4, and statistically significant.[11]

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, in columns (2) to (4) of Table OA.4, we conduct tests using other potential event dates, such as the State Senate and Assembly votes (see Section 2). All coefficients on the California-headquarter indicator variable are insignificant for alternative event dates.

5 Why do shareholders react so strongly to the adoption of the gender quota?

As our results show, the decline in the market value of equity at both Californian and non-Californian firms around the quota announcement is substantial. According to Ahern and Dittmar (2012), a potential source of this value loss could be a deterioration of board quality induced by the quota. Specifically, if not enough qualified female candidates are available, the friction created by a gender quota would be costly for affected firms.

Another reason for the observed value losses at Californian and non-Californian firms may

[11] Our results are also robust if we apply the portfolio sorts approach proposed by Eckbo, Nygaard, and Thorburn (2018) to address concerns that single event date analysis may be biased due to contemporaneous cross-correlation of (abnormal) stock returns (e.g., Brown and Warner, 1985), see Section 2.3 and Table OA.3 in the Online Appendix.

be that the adoption of the gender quota is interpreted as a signal of California's willingness to legislate non-economic values. Several papers show that the uncertainty associated with possible changes in government policy has implications for firm values (e.g., Pástor and Veronesi (2012; 2013); Brogaard and Detzel (2015); Kelly, Pástor, and Veronesi (2016); Koijen, Philipson, and Uhlig (2016)). If the introduction of the gender quota affects shareholders' expectations that legislation targeting non-economic values will be imposed on firms, negative valuation effects may arise for both Californian and non-Californian firms. Next, we examine the extent to which frictions in the managerial labor market, as well as concerns about future legislation of non-economic values, drive the negative announcement returns to the gender quota in California.

5.1 Frictions on the director labor market

The need to search for additional female directors is costly if the pool of female directors is small (Ahern and Dittmar (2012)). However, even a large pool of female directors may be costly to access if female directors are part of different job networks that firms don't have regular access to (Calvó-Armengol and Jackson (2004); Ferreira, Ginglinger, Laguna, and Skalli (2017); Beaman, Keleher, and Magruder (2018)). In this case, firms have to establish costly new recruiting procedures to identify suitable female candidates.

In Figure 3, we examine how the fraction of female directors changes at Californian and non-Californian firms after the introduction of the gender quota in California. Specifically, we compute the change in mean female board representation relative to the adoption of the quota on September 30th, 2018, in percentage points on a daily basis from July 1st, 2018, to March 31st, 2019, using 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). Californian 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 March 2019. Moreover, Californian firms in need of at least two female directors to comply with the quota's requirement (black dashed line) seem to respond more strongly than both other Californian firms and non-California-headquartered control firms that would (hypothetically) need at least two female directors to fulfill the quota (grey dashed line).[12]

To examine whether frictions in the director labor market are the main driver of the negative announcement returns, we re-run our baseline regression from Table 4, column (3), and add a firm-specific characteristic that captures the extent to which a given firm would face frictions from the introduction of the gender quota. While frictions should be most relevant for affected firms in California, they may also affect firms outside of California, if shareholders expect a similar legislation to follow in other states. Indeed, discussions about gender quotas at the senates of other U.S. states show that the outlined proposals are very similar to SB 826.[13] To allow the effect of frictions to differ between Californian and non-Californian firms, we interact the California-headquarter dummy with a variable that captures the extent to which a given firm would face frictions.

In the first specification, which is reported in column (1) of Table 5, we use a variable that counts the number of female directors a firm needs to appoint to fulfill the quota (# missing female directors) as a measure of the frictions faced by the firm. This variable ranges from zero for firms that comply with the quota at adoption announcement to three for firms with an all-male board comprising six or more directors. Results show that the coefficient on the California-headquarter dummy turns insignificant, while the coefficient on the interaction term between the California-headquarter dummy and the number of missing female directors is negative and significant. Thus, Californian firms facing a larger friction, i.e., firms with a larger female director gap on their board, indeed react more strongly to the gender quota announcement. Frictions do not seem to play a role for non-Californian firms, as indicated by the insignificant coefficient on the friction variable itself. We also observe that the constant remains negative, economically large, and statistically significant at the 1% level. Hence, while frictions on the director labor market explain part of the negative announcement returns of Californian firms, a large part of the negative announcement returns of Californian and non-Californian firms remains unexplained.

[12] In Table OA.5, we provide results of a difference-in-differences regression with results discussed in Section 3 of the Online Appendix. These tests mirror what is shown in Figure 3 and indicate that the differences in the increase in female board representation between the treated Californian firms and the non-Californian control firms turn significant already two months after the introduction of the gender quota.

[13] See, for example, Bill NJ S3469 (New Jersey) and Bill SB 5142 – 2019-20 (Washington).

In column (2), we use quota compliance at adoption as an alternative proxy for frictions a firm may encounter when becoming subject to a board gender quota. Specifically, we define a dummy variable equal to one if a firm complies with California's board gender quota at the time of its adoption, and zero otherwise. There are 51 of 455 (11.21%) Californian firms that are already in compliance with the quota at the time of its adoption. Results in column (2) show that the coefficient on the California-headquarter dummy is now positive and significant, while the coefficient on its interaction term with the dummy for quota non-compliance is negative and significant. In economic terms, the announcement returns of non-compliant Californian firms are -3.20% while those of compliant Californian firms are only -0.44%, which is statistically insignificant. Hence, frictions on the director labor market significantly contribute to the negative announcement returns of Californian firms. They do not, however, contribute to the large and negative announcement returns of non-Californian firms – as indicated by the insignificant coefficient on the friction variable itself. Again, the constant is economically large at -2.16% and statistically highly significant, suggesting that frictions on the managerial labor market do not fully explain the negative announcement returns of Californian and non-Californian firms.

In column (3), we include missing female director fixed effects as a proxy for frictions and interact them with the California-headquarter dummy variable. The set of missing female director fixed effects includes a dummy for three missing female directors required to comply with the quota, a dummy for one missing female director, and a dummy for firms that are already in compliance with the quota. The majority of Californian firms lack two female directors for quota compliance (171 of 455 firms, or 37.58%). Hence, this category is omitted from the set of explanatory variables and thus shows in the constant. By including friction fixed effects, we remove unobserved heterogeneity across firms with different numbers of female directors and directly compare Californian firms with three, one, or zero missing directors to the base case, Californian firms with two missing female directors.

The negative and significant coefficient on the California-headquarter dummy in column (3) suggests that Californian firms in need of two female directors to comply with the quota underperform non-Californian control firms with the same number of missing female directors for (hypothetical) quota compliance by 1.27%. While Californian firms that already comply with the quota have 2.35% higher returns than Californian firms with two missing female directors, announcement returns of Californian firms with one or three female directors missing for quota compliance do not differ significantly from the announcement returns of Californian firms with two missing female directors.[14] This finding suggests that negative announcement returns of Californian firms do not increase linearly in the number of female directors that need to be appointed for quota compliance. This, however, is not conclusive evidence against a friction-based explanation of the negative announcement return to the quota: There may be economies of scale when searching for multiple additional female directors.

Most importantly, results in column (3) show a negative and significant constant of -1.53%, indicating that a large part of the negative announcement returns for both, Californian and non-Californian firms, remains unexplained by our measures of frictions in the director labor market.

5.2 Legislating non-economic values

In this section, we test an alternative explanation for the value losses observed at Californian and non-Californian firms in response to the adoption of the gender quota in California. Specifically, we hypothesize that the quota signals California's willingness to legislate non-economic values. In addition, California has a reputation for being a pioneer in introducing new legislation that other states follow. As stated by SEC commissioner Hester Peirce in her speech at the 2018 Annual SEC conference: "Nothing that happens in California stays in California."[15] Thus, after the introduction of the gender quota in California, shareholders may expect that (i) California will legislate other non-economic values in the future, and (ii) that similar legislation will be introduced in other U.S. states.[16]

[14] Note that, as in column (2), Californian firms already in compliance with the quota at adoption show only a small and insignificant announcement return to the quota's adoption of -0.44% (= -1.53% - 1.27% + 2.35%; t-statistic of the combined effect: -0.75).

[15] Peirce, H.M., 2018, My beef with stakeholders: Remarks at the 17th Annual SEC conference, Center for Corporate Reporting and Governance, available at https://www.sec.gov/news/speech/speech-peirce-092118.

[16] See, for example, Tood S. Purdum, *The Nation: Golden Rules; As California Goes, so Goes the Country?*, The New York Times, Sept. 21, 2003, Kate Conger and Noam Scheiber, *California's Contractor Law Stirs Confusion Beyond the Gig Economy*, The New York Times, Sept. 11, 2019, and Christine Mai-Duc and Lauren Weber, *It Isn't Just Uber: California Prepares for New Gig Worker Rules... and Confusion*, The Wall Street Journal, Dec. 17, 2019.

If the observed value losses following the quota's adoption are due to shareholders updating their expectations that further non-economic values will be imposed on firms, we should observe that firms that are more vulnerable to regulatory interventions react more strongly to the quota's adoption. To identify firms that are particularly vulnerable to changes in the regulatory environment, we use the Economic Policy Uncertainty (EPU) index of Baker, Bloom, and Davis (2016) and follow a procedure similar to that in Koijen, Philipson, and Uhlig (2016) and Akey and Lewellen (2017). Specifically, we estimate the same market model regression used in the calculation of the daily ARs and add the daily change of the EPU index as a second explanatory variable. The coefficient estimate on the daily change in the EPU index provides a measure of the firm's sensitivity to changes in the regulatory environment that is not reflected in the market return. Following Akey and Lewellen (2017), we classify a firm as policy-sensitive if the coefficient on the change in the daily EPU index is significant (i.e., has a p-value below 0.1). According to this definition, in our matched sample comprising 1,232 firms, 136 firms (11.04%) are policy-sensitive, out of the which 51 (37.50%) firms are headquartered in California.

Results from estimating regressions similar to those in Table 5, but augmented with the policy sensitivity measure and its interaction with the California-headquarter dummy variable, are reported in Table 6. The regression reported in column (1) includes the California-headquarter dummy, the policy sensitivity dummy, and their interaction. In columns (2) to (4), we account for the effect of frictions on the director labor market by additionally including the number of missing female directors for quota compliance and its interactions with the California-headquarter dummy (column (2)), the quota compliance dummy and its interactions with the California-headquarter dummy (column (3)), or the missing female director fixed effects interacted with the California-headquarter dummy (column (4)).

In column (1), we find negative and significant coefficients of -1.55% on the constant, -0.51% on the California-headquarter dummy, and -1.86% on the interaction term between the California-headquarter and policy-sensitivity dummies. Hence, while Californian firms underperform non-Californian firms by -0.51%, policy-sensitive Californian firms underperform other non-Californian firms by an economically meaningful -1.86%.

In column (2), we augment the regression with the number of missing female directors for quota compliance and an interaction term between the number of missing female directors and the California-headquarter dummy to control for frictions on the director labor market imposed by the quota's adoption. The coefficient on the interaction term between the California-headquarter dummy and the dummy equal to one for policy-sensitive firms remains negative and significant suggesting that Californian firms that are sensitive to economic policy uncertainty experience more negative announcement returns than other Californian firms. However, the negative and significant coefficient on the California-headquarter dummy is now absorbed by the negative and significant interaction term between the California-headquarter dummy and the variable measuring the number of missing female directors for quota compliance. This suggests that both channels, frictions on the director labor market and anticipation of impending value-destroying regulation, matter to Californian firms. Hence, these results are consistent with fears of impending regulation driving part of the negative announcement returns of Californian firms.[17]

Results in columns (3) and (4) confirm these findings: The coefficient on the interaction term between the California-headquarter dummy and the dummy equal to one for policy-sensitive firms remains negative and significant across both specifications. Results on the California-headquarter dummy, the friction measures, and the interaction terms between the two are virtually identical to those reported in Table 5.

In summary, the results in Table 6 provide evidence in support of frictions driving part of the negative announcement returns of Californian firms to the quota's adoption. Moreover, shareholders' expectations regarding future legislation of non-economic values also contribute to the negative announcement returns: Californian firms that historically showed a sensitivity to policy uncertainty experience a more negative announcement return than less policy-sensitive firms. Hence, our results so far suggest that both the friction-based and the legislation-based

[17]One concern with these results is that the methodology used to classify firms as policy-sensitive captures sensitivity to uncertainty more broadly and not necessarily to economic policy uncertainty. We follow Baker, Bloom, and Davis (2016) and others in mitigating this concern. Specifically, we augment the market model regression used to gauge firms' policy sensitivity by both the daily change in EPU and the daily change in the option implied volatility of the S&P500 (VIX). The results remain virtually identical, suggesting that our model indeed picks up the stock return sensitivity to changes in the regulatory environment of the firm rather than the sensitivity to general changes in uncertainty.

explanations contribute to the observed negative announcement returns at Californian firms.

6 Spillover effects to non-Californian firms

Results presented so far suggest that both frictions on the director labor market and an anticipation of impending regulation of non-economic values contribute to negative announcement returns observed around the California board gender quota's adoption. However, the stock price reaction of Californian firms to the quota adoption captures investors' response to the *adoption* of the board gender quota as well as the response to *expectations of the adoption* of other laws in the future. In the remainder of the paper, we thus restrict the sample to all matched non-California-headquartered control firms and conduct spillover tests based on the propensity of states to follow California's legislative lead. The major benefit of such spillover tests is that non-Californian firms are not directly affected by the quota's adoption. Hence, frictions associated with a gender quota and value impairments resulting from the regulation of non-economic values affect company values only through expectations.

If the introduction of a mandatory gender quota in California raises concerns that other states follow by either introducing gender quotas or other value reducing regulation more generally, firms headquartered in states that are more likely to follow California and pass such laws should react more negatively to the introduction of a gender quota in California. The negative reaction could be driven by concerns that a given state would follow up on California's lead and introduce a gender quota which may be costly for firms to fulfill. It could also be driven by more general concerns that non-economic values may be legislated in a given state and lead to additional costs for firms. In the following, we first examine whether announcement returns are more negative in states that are more likely to follow California's legislation (section 6.1). In the next step, we try to disentangle whether these spillover effects are driven by expectations regarding gender quotas specifically, or legislation of non-economic values more generally (section 6.2).

6.1 As California goes, so goes the nation?

Legislators in several states proposed board gender quotas similar to SB 826 – however, none of them have passed at the time of this writing. These states are Illinois, Massachusetts, New Jersey, New York, and Washington.[18] As our first proxy for the likelihood to follow California, we define a dummy variable, Impending quota (d), equal to one for firms headquartered in one of these states. This variable serves as our first proxy for the propensity that a state follows California in either introducing a gender quota or other value-impairing regulation. Results are reported in column (1) of Table 7. As expected, the coefficient on Impending quota (d) is negative and significant, suggesting that firms headquartered in states that are considering the introduction of a board gender quota react more negatively to the passing of Bill 826 in California.

Our next two proxies for a state's likelihood to follow California's legislation are regulatory indices provided by two libertarian think tanks, the John Locke Foundation and the Cato Institute. We would expect states with an already high regulatory density are more likely to follow California's legislative lead and introduce further (value-reducing) regulation in the future. The John Locke Foundation uses 27 provision in six different categories (Land Use, Labor Market, Utilities, Occupations, Tort, and Insurance Regulations) to compute a "regulatory freedom ranking" for each U.S. state. The last available issue of this ranking is from 2015. Our second measure of states' regulatory environments is developed by the Cato Institute. Their index was issued in 2018 and is based on 2016 data. The last version of their regulatory index encompasses 50 provisions in seven categories (Labor Regulation, Health Insurance, Occupational Licensing, Eminent Domain, Liability System, Land and Environment Regulation, and Utility Deregulation). We use the John Locke Foundation ranking and the Cato index to compute two regulatory score variables. To this end, we group all 50 states into quintiles according to their regulatory ranking and index values, respectively, and assign a score of one to firms headquartered in states with the

[18] See, for example, Anastasia Boden, *Setting quotas on women in the boardroom is probably unconstitutional. It also doesn't work*, Los Angeles Times, Jul. 8, 2019, and Laura Weiss, *California board diversity mandate spreads to other states, Washington*, Roll Call, Jul. 19, 2019.

least restrictive regulation and five for firms in states with the strictest regulation.[19] In columns (2) and (3) of Table 7, we show that firms with headquarters outside California but in states with a higher regulatory density experience more negative announcement returns around the adoption of the gender quota in California, but only the coefficient on the regulatory score provided by the John Locke Foundation is significant at conventional levels.

Our fourth proxy for a state's likelihood to impose legislation of non-economic values on firms is a state's political orientation. Given the opposing views of the Democratic and Republican parties on regulatory issues more generally, we expect Democratic states to be more likely to adopt potentially value-impairing regulation. We collect state-level results of the 2016 Presidential Election for the Democratic Party.[20] In column (4), we show that firms with headquarters outside of California but in states with a higher share of votes obtained by the Democratic Party experience more negative announcement returns around the adoption of the gender quota in California.

Our final two proxies are inspired by the idea that California has often been a leader in introducing new regulation, which is subsequently adopted by other states. Specifically, we hypothesize that states that followed California in the past are the states most likely to follow California in adopting either a gender quota or other regulation imposing non-economic values. To this end, we define two proxies that are meant to capture the extent to which states have followed California's legislation in the past. First, we look at states that followed California in introducing minimum wage laws. Such state laws aim at raising the minimum wage above the currently applicable federal rate of $7.25 per hour. California has been a leader in the expansion of minimum wage laws. On April 4th, 2015, California's Governor Jerry Brown signed Senate Bill 3 into law, which prescribes a minimum wage of $12 per hour, including staged increases to $15 per hour by January 1, 2022. In the wake of the California minimum wage law, several states have moved to raise their state minimum wage above the federal minimum as well. Ten states

[19] For an overview of these state-level regulatory scores, see Panels A and B of Figure OA.1 in the Online Appendix.

[20] For an overview of the state-level results of the Democratic Party during the 2016 Presidential election, see Panel C of Figure OA.1 in the Online Appendix.

raised their minimum wage in 2019, and eleven states raised their minimum wage in 2018.[21] Hence, a minimum wage at the state level that exceeds the federal minimum wage of $7.25 per hour may be not only an indicator of a higher probability of following California in adopting progressive laws, but also a signal that this state is willing to impose costs on companies to fulfill non-economic goals. We define a state's propensity to follow California's legislative lead as the difference in dollars between the state-level minimum wage per hour and the federal minimum wage per hour.[22] Results in column (5) of Table 7 show that firms headquartered in states with a higher gap between the state-level and federal minimum wages react significantly more negatively to the introduction of the quota in California.

Finally, we examine whether announcement returns are stronger for firms headquartered in states that have followed California's cannabis legalization policies. In 1996, California was the first state to legalize the use of cannabis for medical purposes when voters approved Proposition 215 – the Compassionate Use Act of 1996. At the time of the adoption of the gender quota in California, the use of cannabis for medicinal purposes was legal in 32 states and the District of Columbia. In 2016, California voters approved the Adult Use of Marijuana Act through Proposition 64, which legalized cannabis for recreational use. Ten jurisdictions followed.[23] For each state, we compute a Cannabis legalization score, which is a count variable equal to zero if any type of cannabis use is considered illegal. It is 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 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 all use, including for medical applications. We conjecture that states that are similar to California in drug regulation are also more likely to

[21] As of September 30, 2018, 27 states and the D.C. have minimum wages per hour in excess of the federal minimum wage. The highest minimum wages per hour are set by the D.C. ($13.25), California ($12), Massachusetts ($12), and Washington ($12). For an overview of the minimum wage policies of all U.S. states, see Panel D of Figure OA.1 in the Online Appendix.

[22] The number of observations is reduced because we disregard firms headquartered in states that do not define a minimum wage (Alabama, Louisiana, Mississippi, New Hampshire, South Carolina, and Tennessee). In alternative tests, we assume the federal minimum wage of $7.25 for these states and find slightly stronger results.

[23] For an overview of the cannabis policy of all U.S. states, see Panel E of Figure OA.1 in the Online Appendix. The years in brackets indicate the year in which cannabis was legalized for medical use.

follow California's lead in introducing a mandatory gender quota. Thus, firms headquartered in these states may react more negatively to the announcement of the California gender quota. Results are reported in column (6) of Table 7. The coefficient on the cannabis-legalization score is negative and statistically significant at the 10% level, suggesting that firms headquartered in states that are more likely to follow California in the adoption of future legislation do indeed have significantly lower announcement returns to the gender quota in California.

In summary, the results in this section suggest that firms headquartered in states with a high propensity to follow California's lead in legislation experience lower returns in response to California's quota adoption. Note that these results are consistent with an anticipation of frictions due to an impending gender quota or fears of value-reducing legislation more generally. In subsequent tests, we attempt to differentiate between these alternative explanations.

6.2 Non-Californian firms: Frictions vs. legislating non-economic values

Next, we decompose the negative announcement returns to non-Californian firms into the part that is due to concerns of impending legislation of non-economic values and the part resulting from frictions associated with the board gender quota.[24] To this end, we estimate regressions similar to those in Table 7. The dependent variable is two-day CARs at quota adoption and the main independent variables are the six alternative measures of a state's propensity to follow California's legislative lead from Table 7.

To discriminate between the return effect that results from impending legislation of non-economic values and the return effect that results from frictions on the director labor market, we additionally include two interaction terms, one between the proxy variable for the probability to follow California and the dummy for policy-sensitive firms and one between the proxy variable

[24]Note that in our spillover setting, there is an alternative channel through which frictions, such as search costs and a deterioration of board quality imposed by the gender quota, may affect our results: Californian firms may attempt to attract female directors from the boards of non-Californian 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. To control for a potential effect resulting from expectations of female directors being hired away from the board of control sample firms by Californian firms, we alternatively augment the regressions by fixed effects for the number of female directors on the board (ranging from zero to six in our sample). Including these fixed effects leaves our results unchanged.

for the probability to follow California and the number of missing female directors. These interactions allow for a differential impact of the quota announcement on policy-sensitive firms and on firms that face larger frictions if a similar gender quota was introduced in a given state. A negative and significant loading on the interaction term between the proxy variable for the probability to follow California and the dummy for policy-sensitive firms would indicate that a significant part of the negative spillover effect results from concerns of impending value-reducing regulation in states that are more likely to follow California's lead. In contrast, a negative and significant interaction term between the proxy variable for the probability to follow California and the number of missing female directors variable suggests that a significant part of the spillover effects is due to concerns of frictions in the director labor market that would result from these states also adopting a board gender quota.

Results are reported in Table 8. Consistent with our expectations, we find five of the six interaction terms between the proxy variable for the probability to follow California and the dummy for policy-sensitive firms to be negative and significant at the 10% level or better. In contrast, the coefficient on the # missing female directors variable and its interaction with the proxies for the propensity to add value-reducing regulations are insignificant across all six columns, suggesting that the negative spillover effects are likely to be due to more general expectations regarding future legislation of non-economic values, rather than a more specific expectation that a gender quota may be introduced.[25]

A concern with our results could be that other factors, that are correlated with policy sensitivity are responsible for the significant relation between these two measures and the announcement returns. The most obvious candidate is firm size, as larger firms may be more policy-sensitive (Akey and Lewellen, 2017). As explained above (and shown for example in Figure 1), larger firms also have more female directors on average. Moreover, they are more likely to have access to national and international director labor markets and thus may be expected to encounter less frictions when facing a board gender quota. Hence, in a robustness test reported in Table OA.6,

[25]Results from replacing the number of missing female director fixed effects by either the quota compliance dummy variable or missing female director fixed effects as a proxy for frictions on the director labor market are similar to the results reported in Table 8.

we re-estimate Table 8 and augment all six regressions with the natural logarithm of total assets as a measure of firm size and its interaction with the proxy variable for the probability that a firm's headquarter state follows California. The table shows that our results remain very similar, while firm size and its interaction term are insignificant across all six specifications, suggesting that firm size does not drive our results.

7 Discussion and Conclusion

Several studies show that stock prices of companies suffer significant losses when women are added to the board. But most of them struggle explaining what drives the large value losses associated with a quota. For example, Ahern and Dittmar (2012) show significantly negative announcement returns of Norwegian firms to the introduction of a gender quota, arguing that frictions on the director labor market caused the significant drop in the stock prices. This reasoning has recently been questioned by Eckbo, Nygaard and Thorburn (2019), who argue that the pool of qualified female directors was large enough to avoid significant shareholder-borne costs. Similarly, Dobbin and Jung (2011) provide evidence that institutional investors sell their shares after a female director is appointed to the board. As they do not find that the addition of female directors negatively affects firms' operating performance, they conclude that these investors are affected by gender bias. As a result of these studies, in 2017, some observers even asked the provocative question "Is the stock market sexist?"[26]

In this paper, we document that the introduction of a gender quota in California led to similar stock market reactions as in other countries: We find robust and significantly negative announcement returns of Californian firms. Interestingly, however, we also find that stock prices of non-Californian firms in states that are likely to follow California's legislative lead react negatively, even though these firms are not subject to the quota.

Consistent with the notion that there is no direct impact of the Californian gender quota on non-Californian firms in these states, we find that frictions explain part of the announcement re-

[26]See Katie Gilbert, *Is the stock market sexist?*, Pacific Standard, Jan. 25, 2016, available at https://psmag.com/economics/is-the-stock-market-sexist.

turns of Californian firms, but they do not explain the negative spillover effects on non-Californian firms. The negative spillover effects must thus be driven by something else. We propose that shareholders take the introduction of a gender quota as a signal that future legislation of non-economic values will follow and find evidence consistent with this view, i.e., policy-sensitive firms in California and in states that are likely to follow California's legislative lead react strongest to the quota announcement. Our results show that even if the direct costs associated with the quota itself are negligible or non-existent, large negative stock price reactions can occur. Shareholders may simply not appreciate that the company is forced to change its organizational structure to achieve non-monetary goals. Thus, even a cost-neutral policy may have adverse effects on stock prices, if expected costs arise from a perceived shift towards stakeholder value maximization. In times of climate change and a stronger ESG orientation of firms, these expectations may be purely rational and are not necessarily driven by gender bias, as the latter would not be consistent with policy-sensitive firms reacting stronger to a gender quota than other firms.

References

Adams, R.B., 2016, Women on boards: The superheroes of tomorrow?, *Leadership Quarterly* 27, 371-386.

Adams, R.B., and D. Ferreira, 2007, A theory of friendly boards, *Journal of Finance* 62, 217-250.

Adams, R.B., and D. Ferreira, 2009, Women in the boardroom and their impact on governance and performance, *Journal of Financial Economics* 94, 291-309.

Ahern, K.R., and A.K. Dittmar, 2012, The changing of the boards: The impact on firm valuation of mandated female board representation, *Quarterly Journal of Economics* 127, 137-197.

Akey, P., and S. Lewellen, 2017, Policy uncertainty, political capital, and firm risk-taking, Working Paper, University of Toronto.

Baker, S.R., N. Bloom, and S.J. Davis, 2016, Measuring economic policy uncertainty, *Quarterly Journal of Economics* 131, 1593-1636.

Beaman, L., N. Keleher, and J. Magruder, 2018, Do job networks disadvantage women? Evidence from a recruitment experiment in Malawi, *Journal of Labor Economics* 36, 121-157.

Bertrand, M., 2017, The glass ceiling, Becker Friedman Institute for Research in Economics Working Paper No. 2018-38.

Brogaard, J., and A. Detzel, 2015, The asset pricing implications of government economic policy uncertainty, *Management Science* 61, 3-18.

Brown, S.J., and J.B. Warner, 1985, Using daily stock returns: The case of event studies, *Journal of Financial Economics* 14, 3-31.

Calvó-Armengol, A., and M.O. Jackson, 2004, The effects of social networks on employment and inequality, *American Economic Review* 94, 426-454.

Dobbin F., and J. Jung, 2011, Corporate board gender diversity and stock performance: The competence gap or institutional investor bias?, *North Carolina Law Review* 89, 809-838.

Eckbo, B.E., K. Nygaard, and K.S. Thorburn, 2018, Board gender-balancing and firm value, Working Paper, Dartmouth College.

Ferreira, D., Ginglinger, E., Laguna, M.-A., and Y. Skalli, 2017, Board quotas and director-firm matching, Working Paper, London School of Economics.

Greene, D., V.J. Intintoli, and K.M. Kahle, 2019, Do board gender quotas affect firm value? Evidence from California Senate Bill No. 826, *Journal of Corporate Finance*, forthcoming.

Grundfest, J., 2018, Mandating gender diversity in the corporate boardroom: The inevitable failure of California's SB 826, Working Paper, Stanford University Law School.

Hwang, S., A. Shivdasani, and E. Simintzi, 2018, Mandating Women on Boards: Evidence from the United States, Working Paper, Kenan Institute of Private Enterprise.

Kelly, B., L. Pástor, and P. Veronesi, 2016, The price of political uncertainty: Theory and evidence from the option market, *Journal of Finance* 71, 2417-2480.

Koijen, R.S., T.J. Philipson, and H. Uhlig, 2016, Financial health economics, *Econometrica* 84, 195-242.

Matsa, D.A., and A.R. Miller, 2013, A female style in corporate leadership? Evidence from quotas, *American Economic Journal: Applied Economics* 5, 136-169.

Pástor, L., and P. Veronesi, 2012, Uncertainty about government policy and stock prices, *Journal of Finance* 64, 1219-1264.

Pástor, L., and P. Veronesi, 2013, Political uncertainty and risk premia, *Journal of Financial Economics* 110, 229-242.

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

This figure shows female board representation in all listed U.S. firms included in the Russell 3000 index from 2008 to 2018 as well as female board representation in subsets of large firms included in the Russell 1000 index and small firms included in the Russell 2000 index. Index membership is determined as of the annual index constitution date using data from Russell's website. Board data from BoardEx as of the annual index constitution date is used to estimate female board representation.

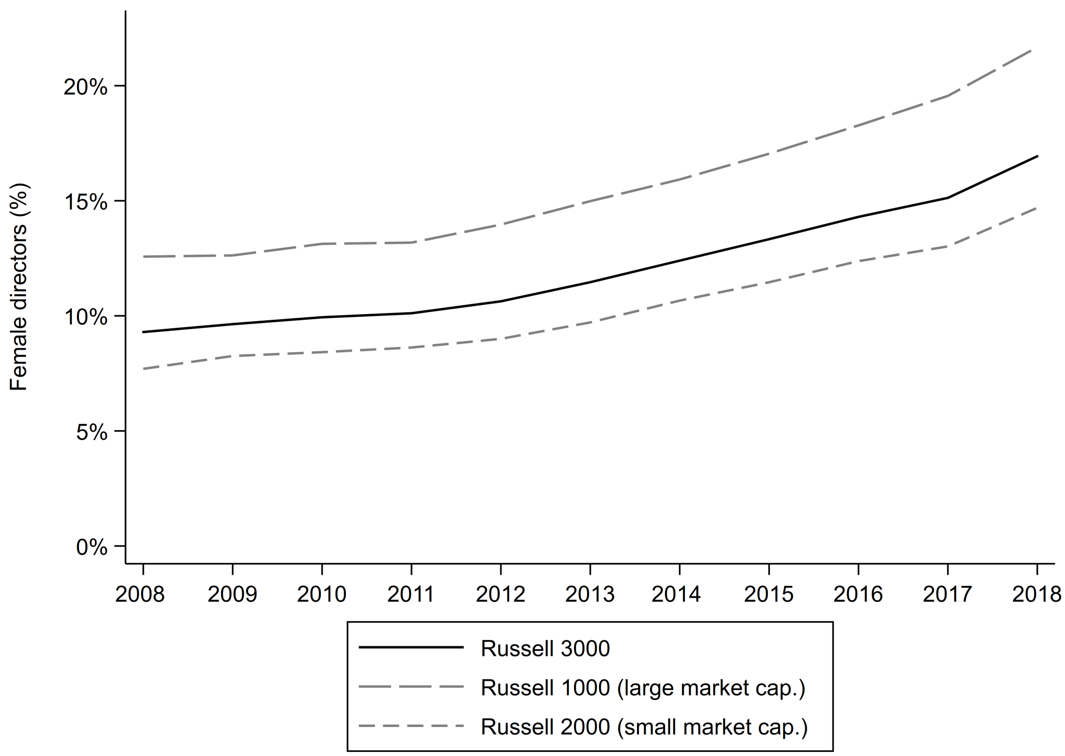

Figure 2: Distribution of newspaper coverage over time

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". The figure displays the weekly fraction of articles that contain these search terms during the time period from December 1, 2017, to November 30, 2018.

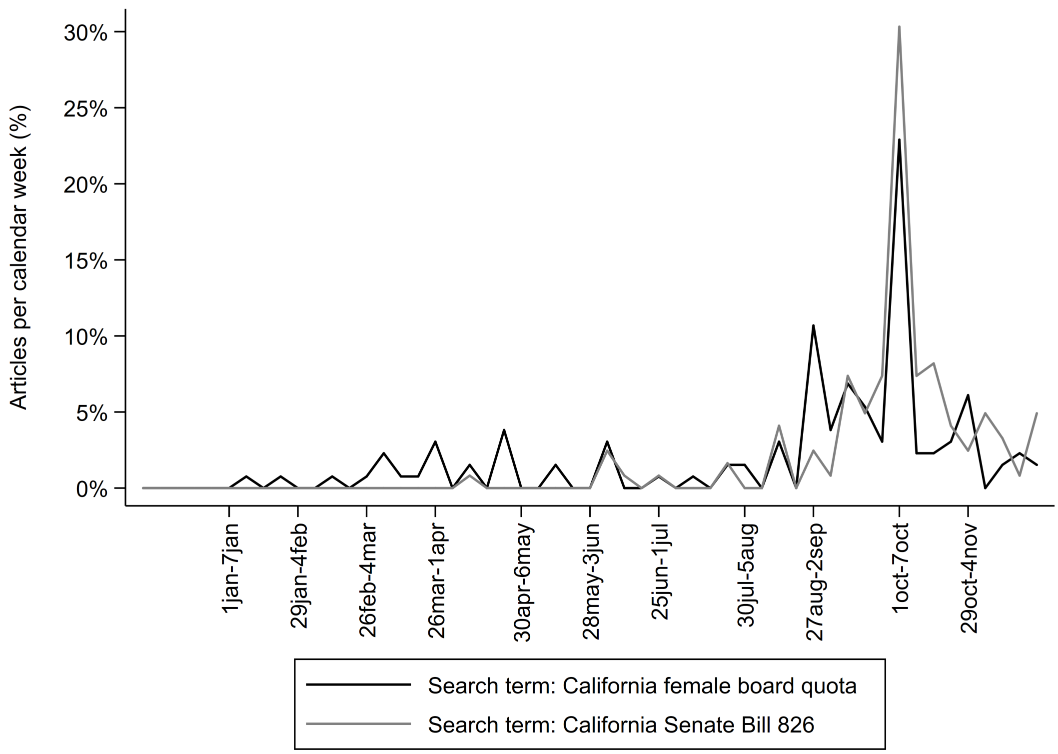

Figure 3: Changes in female board representation around quota adoption
This figure shows daily changes in female board representation for the time period July 1, 2018, to March 31, 2019, relative to the adoption of the quota on September 30, 2018. Solid lines show percentage point changes for a sample 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 firms (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 adoption date. The red vertical line indicates the date the law was signed by the Governor (September 30, 2018). 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 U.S., 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, 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.

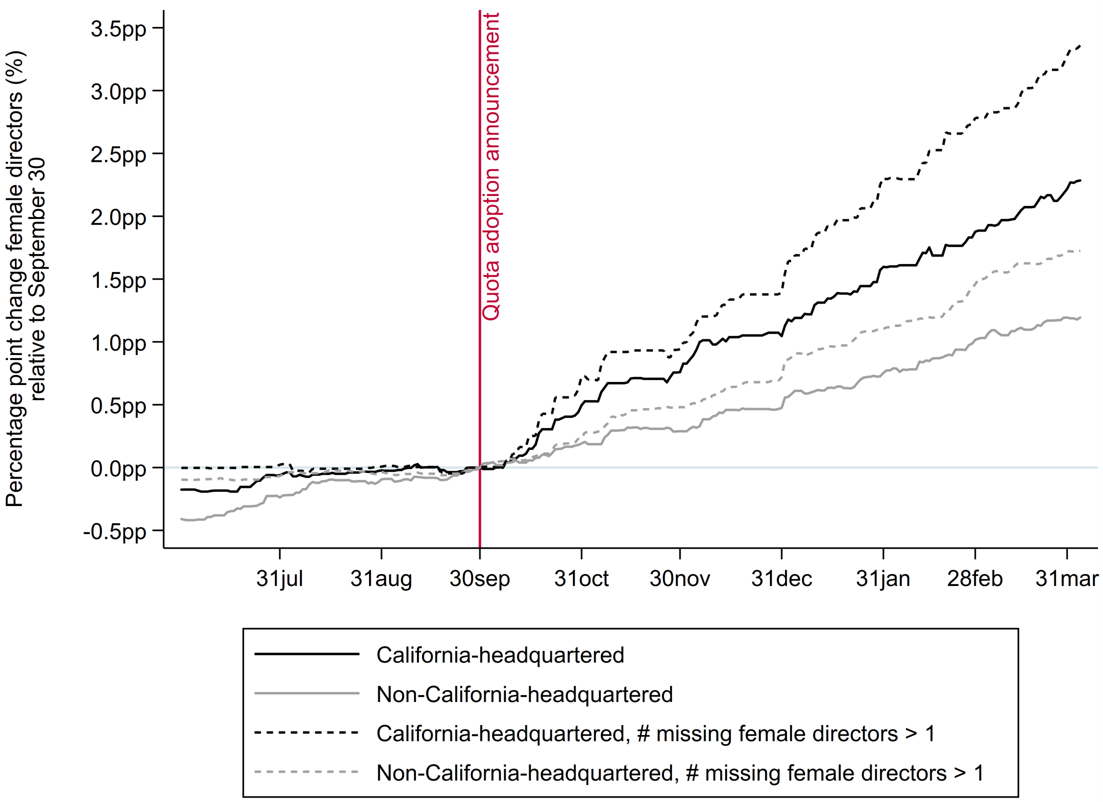

Table 1: Overview of international gender legislation

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 and SB supervisory board ("Aufsichtsrat" in Germany and "Conselho geral e de supervisão" in Portugal).

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 each gender	2010
Iceland	04.03.2010	Listed and private firms >50 employees	BoD	40% each gender; if < 4 directors, one of each gender	2013
France	13.01.2011	Listed firms, private firms >500 employees or >50 million € 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)
India	30.08.2013 / 28.03.2018	Listed firms	BoD	at least 1 woman / 1 independent woman	2015 / 2019 (large), 2020 (small)
Germany	06.03.2015	Listed firms >2,000 employees	SB	30% each gender	2016
Portugal	21.07.2017	Listed firms	BoD, SB	20% each gender / 33.33 % each gender	2018 / 2020

Table 2: Descriptive statistics

Panel A reports descriptive statistics for the sample of firms subject to California's board quota. A firm enters this sample if it is headquartered in California and has a reporting date within one calendar year before the quota's adoption announcement (September 30, 2018) in Compustat. We exclude utility and financial firms (SIC codes 4940-4949 and 6000-6999, respectively), 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, complete return data for the entire five-day event window around the event date (October 1), and availability of director data from BoardEx. Panel B reports descriptive statistics for the matched control sample. To construct this control sample, we draw, for each firm headquartered in California, the three firms closest in size that are active in the same two-digit SIC code industry, are headquartered in another U.S. state or the D.C., and pass the same sample selection criteria as outlined above. 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. Detailed variable definitions are in the Appendix.

Panel A: California-headquartered firms

Firm characteristic	Mean	P25	Median	P75	SD	N
Board size	7.664	6.000	8.000	9.000	1.891	455
Female directors (%)	0.146	0.000	0.143	0.222	0.124	455
# missing female directors	1.686	1.000	2.000	2.000	0.929	455
2021 requ. failed (d)	0.888	1.000	1.000	1.000	0.316	455

Panel B: Industry- and size-matched control firms

Firm characteristic	Mean	P25	Median	P75	SD	N
Board size	7.798	6.000	8.000	9.000	2.030	777
Female directors (%)	0.146	0.000	0.143	0.222	0.124	777
# missing female directors	1.644	1.000	2.000	2.000	0.944	777
2021 requ. failed (d)	0.867	1.000	1.000	1.000	0.339	777

Table 3: Run-up return tests

This table reports differences in cumulative abnormal returns between California-headquartered firms and U.S. non-California-headquartered matched control firms for different event windows that predate the quota adoption announcement. We also report results from t-tests against zero for different cumulative abnormal return measures for the subsample of firms headquartered in California (CA HQ (d) = 1) and for the matched control sample comprising 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 construction of the matched control sample is described in detail in the caption of Table 2. All cumulative abnormal return measures are winsorized at the 1st and 99th percentiles. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix.

	CA HQ (d) = 1			CA HQ (d) = 0			Differences	
	Mean	SE	N	Mean	SE	N	Mean	SE
CAR (-250,-1)	-2.40%***	0.49%	444	-3.23%***	0.36%	751	0.83%	0.60%
CAR (-200,-1)	3.48%***	1.17%	453	1.06%	0.84%	769	2.42%*	1.41%
CAR (-150,-1)	0.54%	1.41%	454	0.94%	0.96%	773	-0.39%	1.65%
CAR (-100,-1)	-3.15%**	1.32%	455	-2.30%**	0.97%	776	-0.85%	1.62%
CAR (-50,-1)	-5.14%***	1.01%	455	-4.78%***	0.73%	776	-0.36%	1.22%

Table 4: Market reaction to the quota's adoption announcement

This table reports differences in abnormal returns to the announcement of the adoption 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)). Across columns, we vary the length of the event window. 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. All abnormal return measures are winsorized at the 1st and 99th percentiles. Heteroscedasticity-consistent standard errors are reported in parentheses. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix.

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**	-0.32*	-0.72***	-0.77**	-1.03***	-1.08***
	(0.19)	(0.18)	(0.26)	(0.31)	(0.33)	(0.39)
Constant	-0.86***	-0.63***	-1.53***	-1.51***	-0.53**	-0.73***
	(0.12)	(0.11)	(0.16)	(0.20)	(0.22)	(0.25)
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

Table 5: Quota compliance and announcement returns

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), variables that capture the extent to which firms comply with the quota at adoption, and interaction terms between the two. The quota compliance measure in Column 1 is a count variable that equals the number of female directors that need to be appointed, based on board size as of the event date, to comply with the requirements of California's board gender quota. The quota compliance measure in Column 2 is a dummy set equal to one if a firm complies with California's board gender quota at adoption and zero otherwise. In Column 3, we interact dummies for the number of female directors that Californian firms need to appoint, based on board size as of the event date, to comply with the requirements of California's board gender 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. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. The construction of the matched sample is described in detail in the caption of Table 2. Heteroscedasticity-consistent standard errors are reported in parentheses. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix.

Dependent variable:	CAR(0,1)		
	(1)	(2)	(3)
CA HQ (d)	0.46	1.72**	-1.27***
	(0.51)	(0.72)	(0.38)
# missing female directors	0.14		
	(0.17)		
CA HQ (d) × # missing female directors	-0.70**		
	(0.29)		
CA HQ (d) × 2021 requ. failed (d)		-2.76***	
		(0.77)	
2021 requ. failed (d)		0.73	
		(0.45)	
# missing female directors = 0 | CA HQ (d) = 1			2.35***
			(0.68)
# missing female directors = 1 | CA HQ (d) = 1			0.69
			(0.47)
# missing female directors = 3 | CA HQ (d) = 1			0.36
			(0.62)
Constant	-1.75***	-2.16***	-1.53***
	(0.30)	(0.41)	(0.16)
F-test # missing female directors FE | CA HQ (d) = 1	-	-	4.09***
R^2	0.01	0.02	0.01
N	1,232	1,232	1,232

Table 6: Policy sensitivity and quota announcement returns

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 proxy for a firm's vulnerability to future regulation, and an interaction term between the latter variable and the California-headquarter dummy. The proxy for a firm's vulnerability to future regulation is a dummy variable set equal to one if a firm's stock returns in the 250-day estimation window depend significantly on changes of the daily Economic Policy Uncertainty index of Baker, Bloom, and Davis (2016), zero otherwise. Regressions reported in Columns 2 to 4 additionally include variables that capture the extent to which firms comply with the quota at adoption and interaction terms between these variables and the California-headquarter dummy. The quota compliance measure in Column 2 is a count variable that equals the number of female directors that need to be appointed, based on board size as of the event date, to comply with the requirements of California's board gender quota. The quota compliance measure in Column 3 is a dummy set equal to one if a firm complies with California's board gender quota at adoption and zero otherwise. In Column 4, we interact dummies for the number of female directors that Californian firms need to appoint, based on board size as of the event date, to comply with the requirements of California's board gender 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. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. The construction of the matched sample is described in detail in the caption of Table 2. Heteroscedasticity-consistent standard errors are reported in parentheses. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix.

Dependent variable:	CAR(0,1)			
	(1)	(2)	(3)	(4)
CA HQ (d)	-0.51*	0.64	1.86**	-1.05***
	(0.28)	(0.51)	(0.73)	(0.39)
CA HQ (d) × Policy sensitive firm (d)	-1.86**	-1.81**	-1.77**	-1.78**
	(0.75)	(0.74)	(0.74)	(0.74)
Policy sensitive firm (d)	0.20	0.17	0.18	0.20
	(0.50)	(0.49)	(0.49)	(0.50)
CA HQ (d) × # missing female directors		-0.69**		
		(0.29)		
# missing female directors		0.13		
		(0.17)		
CA HQ (d) × 2021 requ. failed (d)			-2.70***	
			(0.77)	
2021 requ. failed (d)			0.72	
			(0.45)	
# missing female directors = 0 | CA HQ (d) = 1				2.28***
				(0.68)
# missing female directors = 1 | CA HQ (d) = 1				0.68
				(0.47)
# missing female directors = 3 | CA HQ (d) = 1				0.33
				(0.62)
Constant	-1.55***	-1.76***	-2.18***	-1.55***
	(0.17)	(0.30)	(0.42)	(0.17)
R^2	0.01	0.02	0.02	0.02
N	1,232	1,232	1,232	1,232

Table 7: Spillover tests: The propensity to follow California's legislative lead

This table reports results from pooled ordinary least squares regressions of two-day cumulative abnormal returns (CARs) on state-level variables that serve as a proxy for the likelihood that a firm headquartered in a state becomes subject to a future board gender quota. Across columns, we vary the future quota likelihood proxy across regressions: In Column 1, we use a dummy set equal to one for firms that are headquartered in states that are likely to introduce a board gender quota. In Columns 2 and 3, we use a regulatory score with higher values indicating high regulation as a proxy for a state's willingness to legislate non-economic values on firms. In Column 4, we use state-level measures of political orientation. In Column 5, we use state-level minimum wages as a proxy for a state's willingness to legislate non-economic values on firms. In Column 6, we use a measure that quantifies the extent to which a state has followed California in legislative issues in the past proxied with the similarity in Cannabis legalization. The sample comprises only non-California-headquartered firms. The construction of this sample is described in detail in the caption of Table 2. 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. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. Heteroscedasticity-consistent standard errors are reported in parentheses. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix.

Dependent variable:	CAR(0,1)					
	(1)	(2)	(3)	(4)	(5)	(6)
Impending quota (d)	-0.75**					
	(0.34)					
Regulatory score (JL)		-0.31**				
		(0.12)				
Regulatory score (Cato)			-0.13			
			(0.12)			
% votes Democratic				-3.42*		
				(1.86)		
Excess minimum wage ($)					-0.20**	
					(0.08)	
Cannabis legalization score						-0.25*
						(0.15)
Constant	-1.23***	-0.47	-1.05**	0.20	-1.14***	-1.15***
	(0.21)	(0.43)	(0.47)	(0.94)	(0.23)	(0.28)
R^2	0.01	0.01	0.00	0.00	0.01	0.00
N	777	775	775	777	752	777

Table 8: Frictions vs. legislating non-economic values

This table reports results from pooled ordinary least squares regressions of two-day cumulative abnormal returns (CARs) on state-level variables that serve as a proxy for the likelihood that a firm headquartered in a state becomes subject to future legislation, labeled X, a proxy for a firm's vulnerability to future regulation, the number of missing female directors a firm needs to appoint under the current board size in order to match the requirements postulated by California's board gender quota, as well as interaction terms between the future regulation likelihood proxy (X) and the proxy for a firm's vulnerability to future regulation and the future regulation likelihood proxy (X) and the number of missing female directors. The proxy for a firm's vulnerability to future regulation is a dummy variable set equal to one of a firm's stock returns in the 250-day estimation window depend significantly on changes of the daily Economic Policy Uncertainty Index of Baker, Bloom, and Davis (2016), zero otherwise. Across columns, we vary the future regulation likelihood proxy (X) as indicated above each column: In Column 1, we use a dummy set equal to one for firms that are headquartered in states that are likely to introduce a board gender quota. In Columns 2 and 3, we use a regulatory score with higher values indicating high regulation as a proxy for a state's willingness to legislate non-economic values on firms. In Column 4, we use state-level measures of political orientation. In Column 5, we use state-level minimum wages as a proxy for a state's willingness to legislate non-economic values on firms. In Column 6, we use a measure that quantifies the extent to which a state has followed California in legislative issues in the past proxied with the similarity in Cannabis legalization. The sample comprises only non-California-headquartered firms. The construction of this sample is described in detail in the caption of Table 2. 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. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. Heteroscedasticity-consistent standard errors are reported in parentheses. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix.

Dependent variable:	CAR(0,1)					
X =	Impending quota (d)	Regulatory score (JL)	Regulatory score (Cato)	% votes Democratic	Excess minimum wage ($)	Cannabis legalization score
	(1)	(2)	(3)	(4)	(5)	(6)
X	-0.27	-0.08	0.07	-0.88	-0.27*	-0.51*
	(0.62)	(0.22)	(0.22)	(3.70)	(0.15)	(0.27)
Policy sensitive firm (d)	0.97	1.41	3.50**	7.70***	1.35**	1.29
	(0.63)	(1.41)	(1.54)	(2.97)	(0.65)	(0.82)
X × Policy sensitive firm (d)	-2.31**	-0.38	-0.91**	-14.87***	-0.55**	-0.76*
	(0.92)	(0.40)	(0.38)	(5.58)	(0.24)	(0.46)
# missing female directors	0.18	0.53	0.37	0.48	-0.03	-0.21
	(0.21)	(0.45)	(0.48)	(1.00)	(0.24)	(0.30)
X × # missing female directors	-0.14	-0.11	-0.06	-0.69	0.08	0.21
	(0.35)	(0.13)	(0.13)	(1.99)	(0.09)	(0.16)
Constant	-1.65***	-1.48*	-2.05**	-1.32	-1.24***	-0.93*
	(0.39)	(0.80)	(0.84)	(1.88)	(0.43)	(0.55)
R^2	0.01	0.01	0.01	0.01	0.01	0.01
N	777	775	775	777	752	777

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).

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. 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 125 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
Board size	Number of directors on the board.	BoardEx
Female directors (%)	Fraction of directors on the board that are female.	BoardEx
# missing female directors	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
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
Policy sensitive firm (d)	Dummy variable equal to one if the firm's stock returns in the 250-day estimation window that ends six trading days prior to the event (September 21) depend significantly (p-value $<$ 0.1) on changes (in percent) of the daily Economic Policy Uncertainty (EPU) index of Baker, Bloom, and Davis (2016) when controlling for daily value-weighted market returns, zero otherwise. The daily EPU index relies on the Newsbank database and is computed as a scaled daily number of articles that appeared in around 1,500 U.S. newspapers and include the triple 'uncertainty' or 'uncertain'; 'economic' or 'economy'; and one of the following policy terms: 'congress', 'deficit', 'Federal Reserve', 'legislation', 'regulation' or 'white house' (including variants).	Baker, Bloom, and Davis (2016)
Impending quota (d)	Dummy variable equal to one for firms headquartered in states that are likely to follow California in introducing a board gender quota (Massachusetts, Illinois, New Jersey, New York, and Washington state), zero otherwise.	Newspaper articles
Regulatory score (JL)	Score variable ranging from one to five with low scores indicating little regulation and high scores high regulation. We group all 50 states and the D.C. into quintiles according to their regulatory ranking and assign a score of one to firms headquartered in states with the least restrictive regulation and five for firms in states with the strictest regulation.The ranking is based on 27 provisions in six categories (Land Use, Labor Market, Utilities, Occupations, Tort, and Insurance Regulations).	John Locke Foundation

Regulatory score (Cato)	Score variable ranging from one to five with low scores indicating little regulation and high scores high regulation. We group all 50 states and the D.C. into quintiles according to their regulatory index values and assign a score of one to firms headquartered in states with the least restrictive regulation and five for firms in states with the strictest regulation. The index is based on 50 provisions in seven categories (Land-use Freedom, Health Insurance Freedom, Labor Market Freedom, Lawsuit Freedom, Occupational Freedom, Miscellaneous Regulatory Freedom, and Cable and Telecommunications).	Cato Institute
% votes Democrats	Fraction of votes obtained by Democratic Party in 2016 Presidential Election in the state where a company's headquarter is located.	Politico
Excess minimum wage ($)	Difference in dollars between the state and federal minimum wage per hour, estimated as the state-wide minimum wage per hour less the federal wage of $7.25 per hour.	NCSL
Cannabis legalization score	Score 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 decriminalized, and equal to three if the recreational consumption of cannabis is legal. States that have only legalized the use of Cannabidiol (CBD) oils are coded as zero because most 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

Online Appendix to: As California goes, so goes the nation? Board gender quotas and the legislation of non-economic values

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1 Introduction

The purpose of this Online Appendix is to provide details and results of additional tests briefly mentioned in the paper "As California goes, so goes the nation? Board gender quotas and the legislation of non-economic values". In Section 2, we discuss alternative specifications of our baseline difference-in-differences estimations of the quota announcement effect provided in the paper, including tests that address the potential concern that our results are biased due to contemporaneous cross-correlation of (abnormal) stock returns. Results of these tests are reported in Tables OA.1 to OA.4. Section 3 describes the results from an analysis of changes in female board representation around the adoption of the quota. The results are reported in Table OA.5. The analysis described in Section 4 aims at ruling out concerns that firm size drives the effect of policy sensitivity on the abnormal returns of the non-California-headquartered firms. The results are reported in Table OA.6. Finally, in Section 5, we discuss the spatial distribution of the state-level proxies for the probability to follow California's legislative lead, as displayed in Figure OA.1, and explain how we construct these variables.

2 Alternative difference-in-differences specifications

2.1 Expected return estimation and alternative control sample

In this sub-section, we show that our main difference-in-differences results, reported in Table 4 and discussed in Section 4.1 of the paper, remain unaffected when we use a market-adjusted model to compute daily expected returns in our event study analysis (Panel A of Table OA.1) or when we use all non-Californian firms that pass the sample selection criteria as control firms (Panel B of Table OA.1).

In Panel A of Table OA.1, we report results from re-estimating the regressions in Table 4 of the paper using a market-adjusted model instead of a regression-based market model to predict daily expected returns. In other words, we now compute daily abnormal returns by subtracting, for each firm and each trading day in the event window, the daily market return from the observed daily stock returns. The market-adjusted model assumes a beta of one for all firms and thereby mitigates concerns of misestimated betas. Compared to the results obtained in Table 4 of the paper, the results remain qualitatively unchanged. For instance, the average two-day cumulative abnormal return difference between firms headquartered Californian firms and non-Californian control firms is -0.72%. The constant is -1.50, suggesting that non-Californian firms underperform the market by 1.50%, while Californian firms underperform by -2.22% (column (3)). As in column (3) of Table 4 of the paper, both estimates are significant at the 1% level.

Next, we vary the control group. Recall that in all regressions in the paper, we use a control sample constructed by drawing for each of our 455 treatment firms headquartered in California three non-Californian firms that share the same primary two-digit SIC and are closest in terms of total assets. In the paper, 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, resulting in a sample that comprises 1,232 firms, 455 in the treatment group and 777 in the control group. In Panel B of Table OA.1, we report results from re-estimating our baseline specification reported in Table 4 of the paper using all non-Californian firms that pass our sample selection procedure explained in Section 3.1 of the paper as a control group. The results again remain qualitatively unchanged compared to those reported in Table 4 using the matched control sample.

2.2 Difference-in-differences estimates with treatment- and post dummies

In this sub-section, we conduct a 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. The results obtained when using a four-day event window with two pre-treatment (September 27 and 28) and two post-treatment (October 1 and 2) observations per sample firm are reported in column (1) of Table OA.2. The difference-in-differences estimator, i.e., the coefficient on CA HQ (d) × Post (d), is negative and significant at the 5% level. In terms of economic magnitude, the coefficient estimate suggests a two-day abnormal return of Californian firms that is 0.72% lower than that of non-Californian firms (0.36% per post-treatment day), a number identical to the two-day abnormal return difference of -0.72% reported in column (3) of Table 4 of the paper. Moreover, the coefficient of -0.64% on the Post (d) dummy variable, which is significant at the 1% level, suggests that both Californian and non-Californian firms significantly underperform the market in the two days following the announcement of the quota's adoption. The results in column (2) show that the inclusion of a set of firm-level control variables leaves the results virtually unchanged. In column (3), 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 from column (2), including the treatment dummy, CA HQ (d), as these variables are time-invariant over the four-day sample period used in this analysis. While the significance level is slightly reduced, the estimates remain economically unchanged.

2.3 Calendar-time portfolio analysis

In this sub-section, we conduct robustness tests that address concerns arising from the fact that we study the market reaction of firms to one single event. 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 Californian firms, our treatment sample, and all non-Californian 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:

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

where r_t is the daily equally-weighted return of the portfolio of all Californian (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 Californian and non-Californian firms, we define r_t as the daily difference in portfolio returns of Californian and non-Californian firms. r_{wt} is the daily value-weighted market index return in excess of the daily 1-month U.S. treasury bill rate. As a proxy for the market return, we use the return of a self-computed, value-weighted portfolio 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), 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 OA.3. As in Table 4 of the paper, 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 the paper, we find announcement returns to the introduction of the quota to be significantly more negative for Californian firms, as shown in the last two columns of the table. Also consistent with results reported in Table 4 of the paper, we find non-Californian firms to react negatively to the quota's adoption in California as well. In terms of economic magnitude, the results obtained here are similar to those reported in Table 4 of the paper. For instance, we find the two-day cumulative abnormal return, which includes the event day and the day after (CAR(0,1)), to be -1.44% for non-Californian firms. Californian firms react even more negatively: Their two-day announcement return is -2.16%. The difference between Californian and non-Californian firms of -0.72% is identical to the estimate reported in column (3) of Table 4 in the paper. All these estimates remain statistically significant at the 5% level. Hence, accounting for potential contemporaneous cross-correlation resulting from a single event does not materially affect our results.

2.4 Adding firm-level covariates to the difference-in-differences analysis

In this sub-section, we analyze whether our difference-in-differences results are robust to controlling for a set of firm-level covariates. 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 reported in column (3) of Table 4 in the paper. In column (1) of Table OA.4, we therefore re-estimate our baseline regression augmented with a set of firm specific control variables. The coefficient on the dummy indicating whether a firm is headquartered in California remains economically similar to that reported in Table 4, and is statistically significant at the 5% level.

2.5 Alternative event dates

Our event study results 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 re-estimate our difference-in-differences analysis for other potential quota-related event dates. These are the day of the introduction of the law (January 3) and the day after, the day of the successful Senate vote (May 31) and the day after, and a three-day event window that includes the day of the Assembly vote (August 29), the day of the second Senate vote (August 30), and the day after. Results are reported in columns (2) to (4) of Table OA.4. All three coefficients on the California-headquarter indicator variable are insignificant, suggesting that the market reaction to the California gender quota was confined to the days after the Governor signed the law. The choice of our event window is further justified by the patterns reported in Figure 2 in the paper, which shows the distribution of the newspaper coverage of the gender quota in California.

3 Changes in female board representation around the adoption of the quota

To analyze the pattern displayed in Figure 3 in the paper in a regression framework, we estimate regressions at the firm-month level with the fraction of female directors on the board as dependent variable over the period September 2018 to March 2019. Hence, we obtain a firm-month panel containing up to seven 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 2018, 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. The coefficients on the month dummies indicate the percentage points by which female board representation has changed, on average across all sample firms, in the respective month compared to the introduction of the quota at the end of September 2018. The coefficients on the interaction terms between the month dummy and California-headquarter dummies are the difference-in-differences estimators, that is, the average percentage points difference in the change of female board representation of Californian firms relative to the control firms at the end of 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 are expected to show a positive and significant coefficient.

Results in column (1) of Table OA.5 show that female board representation at Californian firms indeed increased relative to the sample of non-Californian control firms. At the end of October, one month after the quota's introduction, the difference amounts to an insignificant 0.25 percentage points. Two months after the quota's introduction, the difference increases to 0.36 percentage points, which is statistically significant at the 10% level. The difference continues to grow monotonically and amounts to 0.98 percentage points in March 2019, which is statistically significant at the 1% level. Compared to the average annual growth rate of female board representation, which amounts to 0.8% in the Russell 3000 index over the period 2008-2018 (see Figure 1 in the paper), this quota-induced increase in female board representation over a mere six-month time period is therefore 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 adoption 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 3 of the paper, 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 1.17 (1.56) percentage points relative to the control firms six months after the quota's adoption. 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 Californian and non-Californian firms significantly increased female board representation in the months after the adoption of the quota – but as the difference-in-differences estimates show, Californian firms even more so.

The public debate around female board representation often emphasizes the number of firms without any female director on the board to stress the most extreme cases of gender inequality.[1] 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 add additional female directors after the quota's adoption. 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 six difference-in-difference estimators are negative and statistically significant at the 5% level or higher. They suggest that the fraction of Californian firms without a female director on the board has gone down by roughly

[1] See, for instance, Joanna S. Lublin, *Why Breaking Into the Boardroom Is Harder for Women*, The Wall Street Journal, Feb. 2, 2018, and Vanessa Fuhrmans, *The Last All-Male Board on the S&P 500 Is No Longer*, The Wall Street Journal, Jul. 24, 2019.

6% relative to the matched control firms by March-end 2019. 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 Californian 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.

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 six months after adoption of the quota.

4 Controlling for firm size when disentangling frictions and the legislation of non-economic values

In this section, we address the concern that our results of the spillover tests reported in Table 8 of the paper could be driven by omitted firm characteristics that are correlated with policy sensitivity. The most obvious candidate is firm size, as larger firms are more likely to be policy sensitive (Akey and Lewellen, 2017). Moreover, they tend to have more female directors on their boards (see Figure 1 in the paper) and are more likely to have access to national and international director labor markets and thus may be expected to face fewer frictions when becoming subject to a female gender quota. Hence, in Table OA.6, we re-estimate the regressions from Table 8 of the paper augmented with the natural logarithm of total assets as a measure of firm size and interactions between firm size and the proxy variables for the probability that a firm's headquarter state follows California. Results remain very similar to those reported in Table 8 in the paper. Moreover, firm size and its interaction term are insignificant across all six columns, suggesting that firm size does not drive our results.

5 Spatial distribution of state-level proxies for the probability to follow California's lead

In this section, we discuss the spatial distribution of our proxy variables for the states' propensity to follow California's legislative lead. These distributions are graphically displayed in Figure OA.1. Moreover, we explain how we construct these proxy variables.

Panels A and B display the distribution of the regulatory score variables across all 50 U.S. states. The two regulatory scores employed in our paper are based on regulatory indices provided by two libertarian U.S. think tanks, the John Locke Foundation and the Cato Institute. The John Locke Foundation uses 27 provision in six different categories (Land Use, Labor Market, Utilities, Occupations, Tort, and Insurance Regulations) to compute a "regulatory freedom ranking" for each U.S. state. The last available issue of this ranking is from the year 2015. The Cato Institute's index was issued in 2018 and is based on 2016 data. The index encompasses 50 provisions in seven categories (Labor Regulation, Health Insurance, Occupational Licensing, Eminent Domain, Liability System, Land and Environment Regulation, and Utility Deregulation). We use the John Locke Foundation ranking and the Cato index to compute two alternative regulatory score variables. To this end, we group all states and the D.C. into quintiles according to their regulatory index values, and assign a score of one to firms headquartered in states with the least restrictive regulation and five for firms in states with the strictest regulation. Panel A shows the spatial distribution of these scores based on the index using John Locke Foundation's data and Panel B using the Cato Institute's. As expected, California is among the most heavily regulated states in both panels.

Panel C shows the fraction of votes obtained by the Democratic Party during the 2016 Presidential Election. Overall, 21 states and the D.C., predominantly located on the East and West coast, were won by the Democratic Party and 30 states by the Republican Party. With 61.6% of the votes obtained by the Democratic Party, California ranks third behind Hawaii (62.3%) and the D.C. (92.8%).

Panel D provides an overview of the minimum wage policies of all U.S. states and the D.C.

Six states do not set their own minimum wages and simply refer to the federal minimum wage of \$7.25 per hour. Three states even set minimum hourly wages below the federal rate, which implies that the federal rate applies. Fourteen states define the minimum hourly wage to be equal to the federal rate of \$7.25. Twenty-eight states and the D.C. set minimum hourly wages in excess of the federal rate, out of which 14 states set hourly minimum wages between \$7.25 and \$10 and 13 states and the D.C. set hourly minimum wages in excess of \$10. Note that the hourly minimum wage set at the state-level can vary within a state. Oregon, for instance, prescribes a minimum wage per hour of \$10.5 for non-urban counties and of \$12 per hour for the Portland Metropolitan Area. In such cases, we use the lower minimum wage. Moreover, note that we disregard minimum wages set at the municipality or city-level as prevalent, for instance, in Berkeley (\$15 per hour). In the paper, we use the difference between the state-level hourly minimum wage and the federal rate as a proxy for the state's propensity to follow California in legislative issues. As of September-end 2018, California is tied with Massachusetts and Washington for second place with a \$12 minimum wage per hour. The D.C. ranks first with \$13.25 per hour.

Panel E provides an overview of the cannabis policy of all U.S. states and the D.C. as of the board gender quota's adoption date (September 30, 2018). In the paper, we make use an index that captures the extent to which cannabis consumption is legalized in a given state. It is equal to zero if any type of cannabis use is considered illegal (17 states; light grey in Panel F), equal to one if the use of cannabis is legal for medical purposes (13 states; grey), equal to two if the recreational consumption of cannabis is illegal, but has been decriminalized (10 states; dark grey), and equal to three if the recreational consumption of cannabis is legal (10 states and the D.C.; black). Note that states that have only legalized the use of Cannabidiol (CBD) oils are coded as zero because most 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. Also note that while the classification underlying the index follows generally observable patterns in legalizing cannabis consumption at the state level, the sequence holds for all but one state. Specifically, Missouri decriminalized the recreational use of Cannabis in 2014 and only legalized the use of cannabis for medical application as defined above after the adoption of California's board gender quota (in November 2018 through ballot measure). If a state legalized the use of cannabis for medical treatment, the numbers in parentheses indicate the year in which it was legalized. California was the first state to legalize Cannabis for medical purpose (in 1996) and is among those states which legalized Cannabis consumption completely.

References

Akey, P., and S. Lewellen, 2017, Policy uncertainty, political capital, and firm risk-taking, Working Paper, University of Toronto.

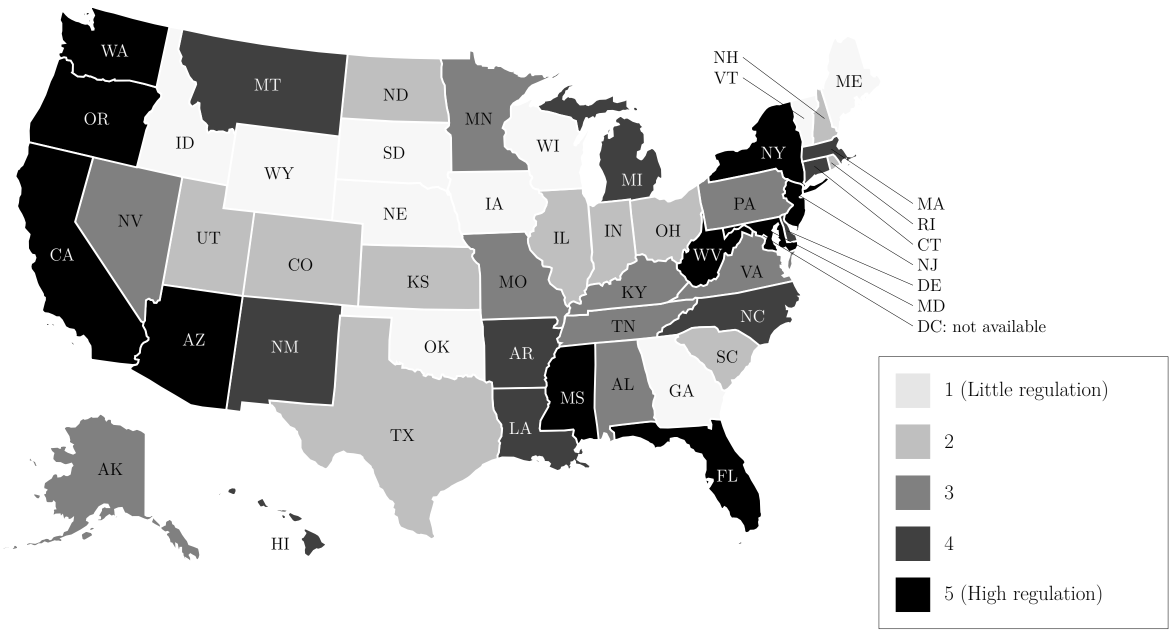
Baker, S.R., N. Bloom, and S.J. Davis, 2016, Measuring economic policy uncertainty, *Quarterly Journal of Economics* 131, 1593-1636.

Eckbo, B.E., K. Nygaard, and K.S. Thorburn, 2018, Board gender-balancing and firm value, Working Paper, Dartmouth College.

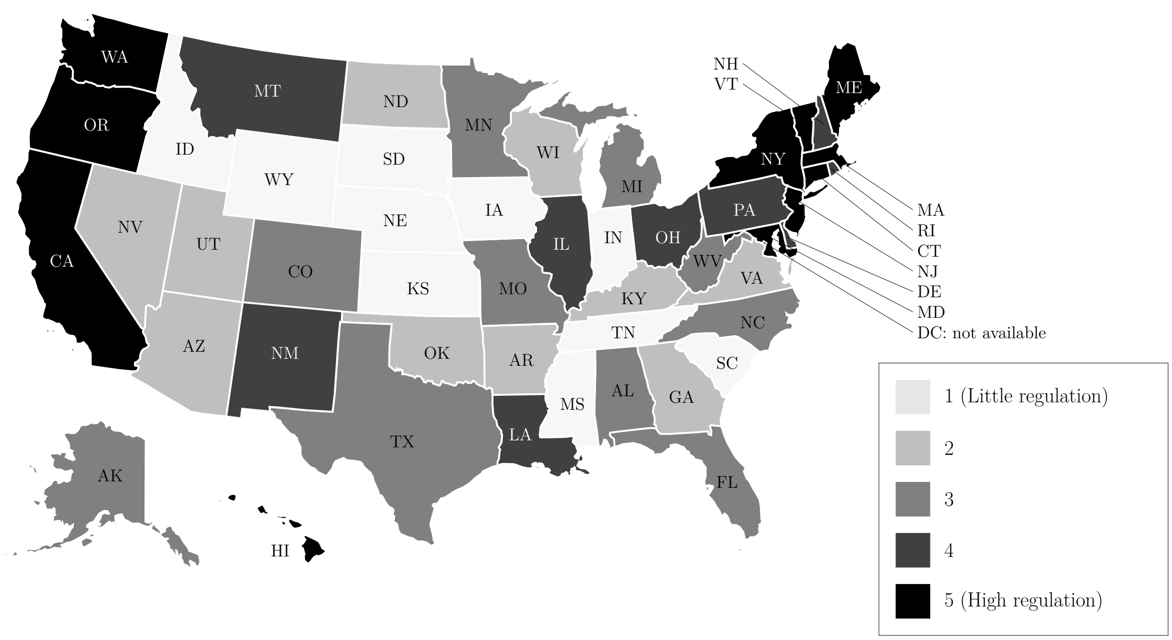
Figure OA.1: Distribution of state-level proxies for the probability to follow California's lead

Panels A and B of this figure show the regulatory scores for each state. Panel C shows the 2016 Presidential Election results for each state and the D.C. Panel D shows the state-level minimum wage per hour for each state and the D.C. Panel E shows the legislation governing cannabis consumption for each state and the D.C. and, if applicable, the year in which cannabis was legalized for medical use. Data sources are provided in the Appendix of the paper.

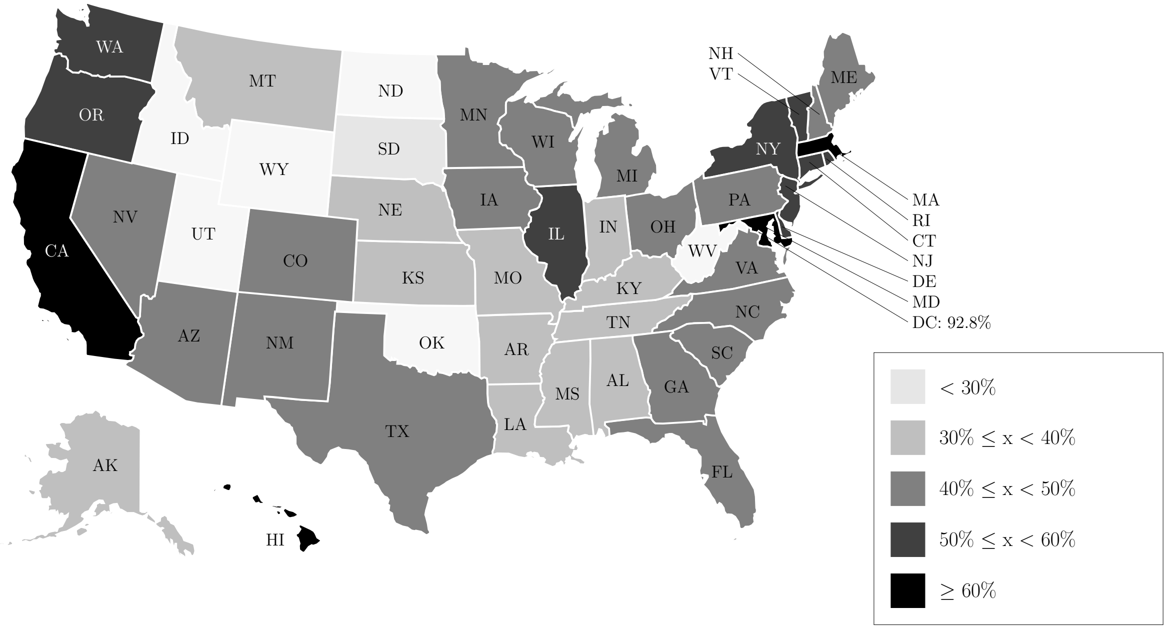
Panel A: Regulatory scores defined using data from the John Locke Foundation



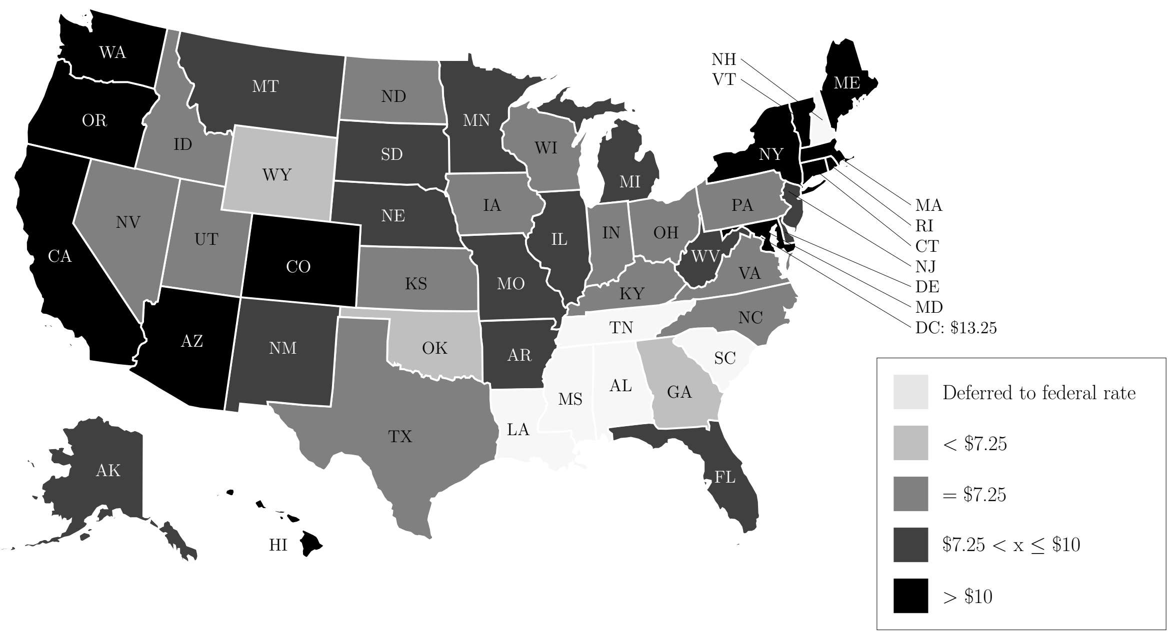
Panel B: Regulatory scores defined using data from the Cato Institute



Panel C: State-level results of Democratic Party during 2016 Presidential Election



Panel D: State-level minimum wages per hour



Panel E: State-level cannabis consumption legislation

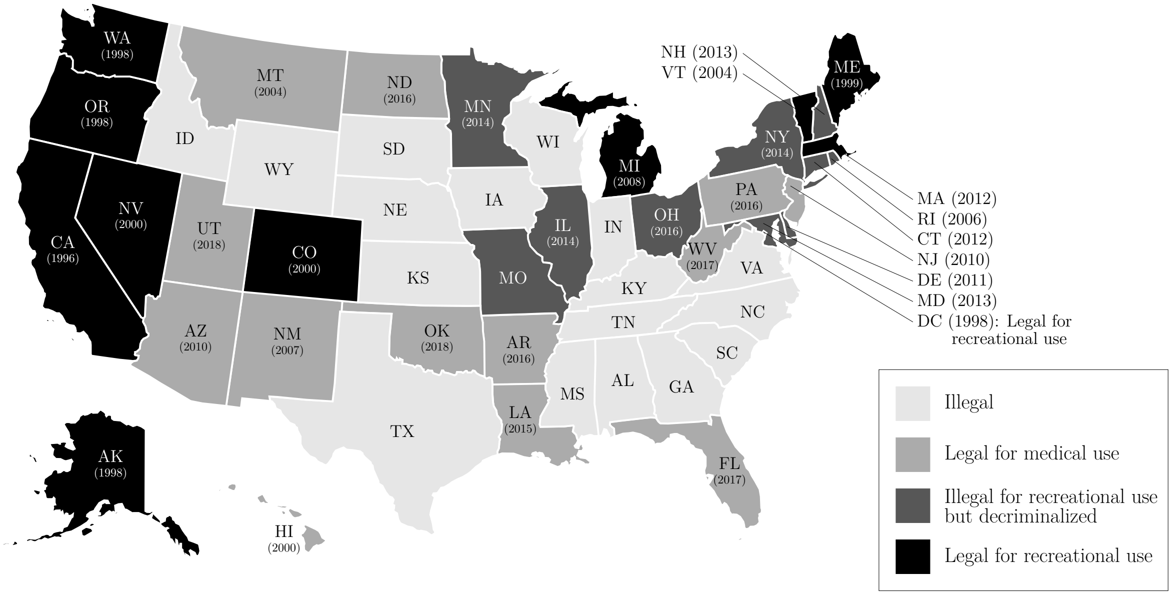


Table OA.1: Robustness tests: Alternative expected return estimations and all non-Californian firms as control firms

This table reports robustness tests of the results shown in Table 3 in the paper. Specifically, it reports differences in abnormal returns around the announcement of the adoption 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)). In Panels A and B, we vary the length of the event window across columns. In our baseline regression shown in Table 3 of the paper, 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. In Panel A of this table, we use a market-adjusted model (instead of a market model) to compute daily expected returns, assuming that the beta of all stocks is equal to one. In our baseline regression shown in Table 3 of the paper, the matched control sample is constructed by drawing, 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. In Panel B of this table, we use all non-California-headquartered firms as control firms that pass the sample selection criteria outlined in Section 3.1 of the paper. All abnormal return measures are winsorized at the 1st and 99th percentiles. Heteroscedasticity-consistent standard errors are reported in parentheses. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix of the paper.

Panel A: 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***
	(0.19)	(0.18)	(0.26)	(0.30)	(0.32)	(0.37)
Constant	-0.89***	-0.58***	-1.50***	-1.47***	-0.50**	-0.75***
	(0.12)	(0.11)	(0.16)	(0.19)	(0.21)	(0.23)
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: Control firms are all non-California-headquartered 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)	-0.52***	-0.42***	-0.90***	-0.97***	-0.98***	-1.04***
	(0.16)	(0.14)	(0.22)	(0.26)	(0.26)	(0.32)
Constant	-0.79***	-0.56***	-1.34***	-1.30***	-0.58***	-0.76***
	(0.06)	(0.06)	(0.09)	(0.10)	(0.11)	(0.13)
R^2	0.01	0.00	0.01	0.01	0.01	0.00
N	2,455	2,455	2,455	2,455	2,455	2,455

Table OA.2: Robustness tests: Difference-in-differences estimations 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 reported in Column 2 additionally includes financial controls while the regression reported in Column 3 additionally includes firm fixed effects. The sample comprises four daily abnormal 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. Daily abnormal returns and all financial ratios are winsorized at the 1st and 99th percentiles. The construction of the matched sample is described in detail in the caption of Table 2 of the paper. All regressions include an intercept, which is not shown for brevity. Standard errors, reported in parentheses, are clustered at the firm level. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix of the paper.

Dependent variable:	AR(t)		
	(1)	(2)	(3)
CA HQ (d) × Post (d)	-0.36**	-0.36**	-0.36*
	(0.18)	(0.18)	(0.20)
Post (d)	-0.64***	-0.64***	-0.64***
	(0.11)	(0.11)	(0.13)
CA HQ (d)	-0.02	0.01	
	(0.12)	(0.12)	
ln(Total assets)		0.02	
		(0.02)	
Leverage		0.12	
		(0.24)	
ROA		-0.18	
		(0.26)	
PPE / TA		0.75**	
		(0.30)	
R&D / TA		-0.37	
		(0.49)	
Firm FE	No	No	Yes
R^2	0.02	0.02	0.27
N	4,928	4,928	4,928
Firms	1,232	1,232	1,232

Table OA.3: Robustness tests: Accounting for the cross-sectional dependence of returns

This table reports cumulative abnormal stock returns for two portfolios, one comprising California-head-quartered 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 for 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. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix of the paper.

	CA HQ (d) = 1			CA HQ (d) = 0			Differences	
	Mean	SE	N	Mean	SE	N	Mean	SE
AR (0)	-1.24%**	0.49%	455	-0.82%**	0.40%	777	-0.42%*	0.24%
AR (1)	-0.92%*	0.49%	455	-0.62%	0.40%	777	-0.30%	0.24%
CAR (0,1)	-2.16%***	0.34%	455	-1.44%**	0.29%	777	-0.72%**	0.17%
CAR (-1,1)	-2.14%**	0.28%	455	-1.39%**	0.23%	777	-0.74%*	0.14%
CAR (0,2)	-1.48%*	0.29%	455	-0.61%	0.24%	777	-0.88%**	0.14%
CAR (-2,2)	-1.63%	0.22%	455	-0.71%	0.18%	777	-0.92%*	0.11%

Table OA.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)) and a set of financial control variables. In Column 1, CARs are estimated over a two-day event window that includes the event date and the first day after (i.e., October 1 and 2). In Column 2, CARs are estimated over a two-day event window that includes the day of the introduction of the law (January 3) and the day after. In Column 3, CARs are estimated over a two-day event window that includes the day of the successful Senate vote (May 31) and the day after. In Column 4, CARs are estimate over a three-day event window that includes 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 six days before the event. As a proxy for the market return, we use the return of a self-computed, value-weighted market index consisting of all sample firms. Cumulative abnormal returns and all financial ratios are winsorized at the 1st and 99th percentiles. The construction of the matched sample is described in detail in the caption of Table 2. All regressions include an intercept, which is not shown for brevity. Heteroscedasticity-consistent standard errors are reported in parentheses. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix of the paper and of this Online Appendix.

Dependent variable:	CAR(0,1)	CAR(Jan. 3, Jan. 4)	CAR(May 31, Jun. 1)	CAR(Aug. 29, Aug. 31)
Event(s):	Law signed by Governor	Law introduced	Successful Senate vote	Successful Assembly vote and second Senate vote
	(1)	(2)	(3)	(4)
CA HQ (d)	-0.64**	0.16	0.26	0.37
	(0.28)	(0.26)	(0.22)	(0.30)
ln(Total assets)	0.02	-0.15**	-0.18***	-0.25***
	(0.07)	(0.08)	(0.06)	(0.08)
Leverage	1.62**	0.76	0.14	-0.76
	(0.75)	(0.68)	(0.61)	(0.77)
ROA	-0.95	0.52	0.51	-1.86**
	(0.74)	(0.81)	(0.77)	(0.83)
PPE / TA	1.93**	-0.05	-0.25	0.09
	(0.84)	(0.81)	(0.56)	(0.76)
R&D / TA	-0.98	2.52*	1.95	-0.08
	(1.46)	(1.50)	(1.33)	(1.45)
R^2	0.02	0.02	0.02	0.05
N	1,232	1,184	1,220	1,229

Table OA.5: Female board representation 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 variable and the month dummy variables. For each California-headquartered firm and the sample of control firms, we compute board characteristics for the end of September (the base month) as well as for the end of October 2018 to March 2019. In Column 1, we use the fraction of directors on the board that are female as the dependent variable. In Column 2, we restrict the sample to firms that require at least one additional female director to fulfill the quota at the quota's adoption date, and in Column 3, we restrict the sample to firms that require at least two additional female directors to fulfill the quota at the quota's adoption date. In Column 4, we use the sample from Column 1 but replace the dependent variable with a dummy variable set equal to one if a firm at the end of a given month has no female director on the board. The construction of the matched sample is described in detail in the caption of Table 2 of the paper. All regressions include firm fixed effects and an intercept, which is not shown for brevity. Standard errors, reported in parentheses, are clustered at the firm level. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix of the paper.

Dependent variable:	Female directors (%)			No female (d)
Sample:	Full	# missing female directors > 0	# missing female directors > 1	Full
	(1)	(2)	(3)	(4)
CA HQ (d) × October-end (d)	0.25	0.34*	0.49**	-0.02**
	(0.16)	(0.17)	(0.23)	(0.01)
CA HQ (d) × November-end (d)	0.36*	0.44**	0.49*	-0.02**
	(0.20)	(0.21)	(0.28)	(0.01)
CA HQ (d) × December-end (d)	0.50**	0.58**	0.69**	-0.03***
	(0.23)	(0.25)	(0.32)	(0.01)
CA HQ (d) × January-end (d)	0.73***	0.86***	1.12***	-0.05***
	(0.27)	(0.29)	(0.39)	(0.01)
CA HQ (d) × February-end (d)	0.77***	0.87***	1.21***	-0.05***
	(0.30)	(0.32)	(0.43)	(0.01)
CA HQ (d) × March-end (d)	0.98***	1.17***	1.56***	-0.06***
	(0.33)	(0.36)	(0.46)	(0.02)
October-end (d)	0.22***	0.21**	0.24**	-0.00
	(0.08)	(0.09)	(0.11)	(0.00)
November-end (d)	0.32***	0.35***	0.49***	-0.01
	(0.11)	(0.11)	(0.14)	(0.00)
December-end (d)	0.49***	0.54***	0.75***	-0.01*
	(0.13)	(0.14)	(0.18)	(0.01)
January-end (d)	0.77***	0.83***	1.20***	-0.02***
	(0.14)	(0.15)	(0.21)	(0.01)
February-end (d)	1.04***	1.16***	1.57***	-0.02***
	(0.17)	(0.18)	(0.23)	(0.01)
March-end (d)	1.20***	1.34***	1.76***	-0.03***
	(0.19)	(0.20)	(0.27)	(0.01)
Firm FE	Yes	Yes	Yes	Yes
R^2	0.05	0.06	0.10	0.03
N	8,520	7,454	4,955	8,520
Firms	1,232	1,075	715	1,232

Table OA.6: Frictions vs. legislating non-economic values: Controlling for firm size
This table reports results from pooled ordinary least squares regressions of two-day cumulative abnormal returns (CARs) on state-level variables that serve as a proxy for the likelihood that a firm headquartered in a state becomes subject to future legislation, labeled X, a proxy for a firm's vulnerability to future regulation, the number of missing female directors a firm needs to appoint under the current board size in order to match the requirements postulated by California's board gender quota, as well as interaction terms between the future regulation likelihood proxy (X) and the proxy for a firm's vulnerability to future regulation and the future regulation likelihood proxy (X) and the number of missing female directors. In contrast to Table 8 of the paper, we add the natural logarithm of total assets and an interaction term between the future regulation likelihood proxy (X) and the natural logarithm of total assets to the model. The proxy for a firm's vulnerability to future regulation is a dummy variable set equal to one of a firm's stock returns in the 250-day estimation window depend significantly on changes of the daily Economic Policy Uncertainty Index of Baker, Bloom, and Davis (2016), zero otherwise. Across columns, we vary the future regulation likelihood proxy (X) as indicated above each column: In Column 1, we use a dummy set equal to one for firms that are headquartered in states that are likely to introduce a board gender quota. In Columns 2 and 3, we use a regulatory score with higher values indicating high regulation as a proxy for a state's willingness to legislate non-economic values on firms. In Column 4, we use state-level measures of political orientation. In Column 5, we use state-level minimum wages as a proxy for a state's willingness to legislate non-economic values on firms. In Column 6, we use a measure that quantifies the extent to which a state has followed California in legislative issues in the past proxied with the degree of Cannabis legalization. The sample comprises only non-California-headquartered firms. Firms in this sample are selected by choosing the three closest control 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. Cumulative abnormal returns are winsorized at the 1st and 99th percentiles. Heteroscedasticity-consistent standard errors are reported in parentheses. *, **, and ***, indicate statistical significance at the 10%, 5%, and 1% level, respectively. Detailed variable definitions are in the Appendix of the paper.

Dependent variable:	CAR(0,1)					
X =	Impending quota (d)	Regulatory score (JL)	Regulatory score (Cato)	% votes Democratic	Excess minimum wage ($)	Cannabis legalization score
	(1)	(2)	(3)	(4)	(5)	(6)
X	-1.13	-0.09	-0.28	-3.02	-0.35	-0.59
	(1.63)	(0.57)	(0.58)	(9.44)	(0.37)	(0.70)
Policy sensitive firm (d)	0.92	1.40	3.55**	7.61**	1.33**	1.27
	(0.63)	(1.40)	(1.52)	(2.97)	(0.65)	(0.82)
X × Policy sensitive firm (d)	-2.26**	-0.38	-0.93**	-14.71***	-0.55**	-0.76*
	(0.92)	(0.40)	(0.37)	(5.58)	(0.24)	(0.46)
# missing female directors	0.13	0.50	0.24	0.41	-0.08	-0.25
	(0.23)	(0.47)	(0.49)	(1.07)	(0.26)	(0.31)
X × # missing female directors	-0.05	-0.12	-0.03	-0.60	0.08	0.21
	(0.38)	(0.13)	(0.13)	(2.14)	(0.09)	(0.16)
ln(Total assets)	-0.08	-0.07	-0.23	-0.20	-0.07	-0.05
	(0.12)	(0.24)	(0.27)	(0.57)	(0.13)	(0.15)
X × ln(Total assets)	0.12	0.00	0.05	0.34	0.01	0.01
	(0.19)	(0.07)	(0.07)	(1.13)	(0.05)	(0.08)
Constant	-1.06	-1.00	-0.42	-0.01	-0.75	-0.52
	(1.04)	(2.04)	(2.18)	(4.71)	(1.12)	(1.29)
R^2	0.01	0.01	0.01	0.01	0.01	0.01
N	777	775	775	777	752	777

Appendix: Variable definitions

This table reports variable definitions and data sources of all variables used in this Online Appendix but not in the paper. Database mnemonics are in italics (if available).

Variable	Definition	Source
Total assets	Total assets (*at*).	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
ROA	Operating income before depreciation scaled by total assets (*oibdp* / *at*), winsorized at the 1% and 99% level.	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