How costly is forced gender-balancing of corporate boards?*

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Abstract

In 2005, Norway became the first country to mandate gender-balanced corporate boards. We hypothesize that a gender quota reduces director CEO experience and increases board independence. Contrary to prior research, our robust performance estimates fail to reject an overall value-neutral effect of the quota, even for firms with all-male boards. We also show that, while boards lost some CEO experience, firms did not increase board size (to retain key male directors) or change legal form (to avoid the quota), and managed to maintain board network power. We conclude that investors and firms alike viewed the quota as a relatively low-cost constraint.

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1 Introduction

In December of 2005, Norway passed a law requiring its public limited liability companies ("Allmenaksjeselskap", henceforth ASA) to have gender-balanced boards (at least 40% of directors from each gender). Firms were given two years to comply, with a penalty of forced liquidation for noncompliance. The new female directors were relatively independent professionals, possibly adding monitoring benefits not available under the previous male-driven director election process. On the other hand, as shown below, the incoming female directors had low CEO experience, potentially weakening board advisory capacity. The main objective of this paper is to examine whether this shock to director independence and experience lowered firm value. We present a battery of tests ranging from short- and long-run market valuation estimates to corporate actions that might help minimize perceived negative quota effects for shareholders and which help validate market-based inferences.

The Norwegian quasi-experiment permits identification of causal effects that are rare in corporate governance research (Adams, Hermalin, and Weisbach, 2010). While other governance shocks, such as the passage of the 2002 Sarbanes-Oxley Act in the US (SOX), also affect director independence, SOX introduced costly internal control structures that confound inferences. In contrast, the quota regulates gender composition only and thus nothing but director independence and experience. Also, the quota was driven by gender politics, which is exogenous to firm performance. Moreover, as the first of its kind, the quota was relatively unanticipated, suggesting that inferences based on market-valuation estimates have power.[1]

Our empirical analysis benefits from access to an exhaustive national corporate registry covering the population of ASA as well as unregulated private limited liability companies ("Aksjeselskap", henceforth AS).[2] This data allows us to implement two difference-in-difference ("diff-in-diff") approaches. The first contrasts regulated ASA with large AS (top 1% by revenue), which were never considered as candidates for quota regulation. The second compares firms with all-male boards prior to the quota with firms facing a lower quota-induced constraint. These diff-in-diff tests control for country-specific changes in board composition over the sample period 1998 through 2013 that are unrelated to the ASA quota constraint.

[1]Norway's quota has since prompted quota legislation in Spain, Belgium, France, Germany, Iceland, Italy, and the Netherlands. In the US, resolutions urging board gender diversity for public companies have recently been passed by the state legislatures of California and Massachusetts.

[2]The corporate legal forms of ASA and AS correspond, respectively, to PLC and Ltd in the UK., S.A. and S.A.R.L. in France, and AG and GmbH in Germany.

The overall conclusion from our evidence and tests is that investors and firms alike viewed the quota constraint as value-neutral, and that shareholder-borne costs were too small for firms to take evasive actions. This conclusion differs from prior studies of Norway's gender quota. Specifically, In particular, Ahern and Dittmar (2012) (on valuation effects) and Bøhren and Staubo (2014) (on corporate legal conversions) report significantly negative quota effects. However, our replication of these studies uncovers objective econometric issues which, when corrected, also supports our conclusion that the quota had a value-neutral effect on ASA listed on the Oslo Stock Exchange (OSE).

We begin the analysis with shareholder actions designed to minimize negative effects of the quota constraint. Surely, if investors and firms anticipated large negative effects—as prior research suggest—one would expect to see evidence of corporate evasive actions. The most straightforward option is to increase board size to make room for females without having to terminate valued male directors. For example, a five-member board (the average board size in the ASA population) could retain all existing male directors by increasing board size to eight. While a larger board is more costly for shareholders (Jensen, 1993; Boone, Field, Karpoff, and Raheja, 2007), the marginal cost of increasing board size effectively places an upper limit on the shareholder-borne costs of the quota constraint.

Our diff-in-diff analysis produces no evidence of a board-size change after quota compliance—not even for ASA with all-male boards. This finding is surprising since the analysis *does* identify a significant decline in director CEO experience for ASA relative to large AS and, in particular, for ASA with all-male boards. It appears that the costs of increasing board size—as perceived by shareholders—exceeded the opportunity cost of lower director CEO experience. This particular interpretation is also supported by our analysis of changes in overall board network power. The diff-in-diff analysis shows that board network power, measured using PageRank scores (Page, Brin, Motwani, and Winograd, 1999), did not change as a result of quota compliance. Thus, overall board *effectiveness*, which is a product of both experience and network power, may have remained largely unchanged despite the decline in director CEO experience.

Our analysis of conversions from ASA to AS is also interesting in this context. While conversion is relatively costly for listed AS (it forces delisting from the OSE), the costs are low for unlisted ASA with no plans to list (it does not affect the option to raise public or private debt, to make private placements of equity, or to trade in the over-the-counter equity market). Firms that convert do not reliably state the conversion reason. Thus, any role played by the quota-induced female director shortfall must be inferred. As shown by As in Bøhren and Staubo (2014) as well, no listed ASA converted over the sample period. However, unlisted ASA regularly convert throughout. We find that the conversion likelihood for these unlisted ASA is statistically unrelated to the shortfall of female directors when one controls for the general time-trend in board gender composition.

Next, we turn to the market valuation of the quota constraint using a combination of short- and long-term stock return analysis. In the short-term event study, we pinpoint multiple quota-related news events for accuracy and power. We also estimate the difference in industry-adjusted abnormal stock returns across portfolios of firms with low- and high female director shortfall. Our event-study tests improve on earlier studies in that they account for the cross-dependency of returns that arises when an event affects all sample firms simultaneously. This cross-dependence is substantial for firms listed on the OSE, and incorporating it into the statistical tests results in insignificant event-induced abnormal returns—also when estimated as in Ahern and Dittmar (2012).

Finally, our analysis fails to produce statistically significant long-term valuation effects, whether we use buy-and-hold stock portfolios, the type of IV-tests for changes in Tobin's Q used by Ahern and Dittmar (2012), or operating performance as in Matsa and Miller (2013). The absence of both short- and long-run valuation effects is, of course, consistent with our finding that shareholders did not find it worthwhile to engage in low-cost actions such as increasing board size or converting from ASA to AS to avoid the quota. Thus, it appears that neither investors nor the firms themselves viewed forced gender balancing as particularly costly, despite a relative decline in CEO experience of regulated boards.

The rest of the paper is organized as follows. Section 2 explains the quota and its legislative time line. Section 3 examines the extent of corporate actions to minimize perceived costs of the quota. Section 4 estimates the market reaction to quota-related events, while Section 5 presents a long-run stock and accounting performance analysis. Section 6 concludes the paper.

2 Quota legislative time-line and data sources

2.1 Legislative time-line

The Norwegian gender quota was signed into corporate law in December of 2005. While Norway's codetermination laws grant employees rights to elect their own directors, the quota law regulates ASA-directors elected by shareholders. Table 1 shows how the mandated fraction of women and men varies with the shareholder-elected board size. For example, in boards with three shareholder-elected directors, at least one (33%) must be female and one male. In boards with four or more directors, the mandated percent female (and male) directors ranges from 50% for a four-member board to 38% for an eight-member board.

Table 2 shows the dates of five major news events that increased the probability of a quota amendment. The events began with February 22, 2002 (event 1), when the Minister of Industry and Trade in an interview with a daily tabloid *Verdens Gang* first signaled political support for a mandatory gender quota. It was a surprising and controversial move, particularly within his own conservative party. Perhaps because of this, he publicly retracted his support the very next day, through an interview in the major national business newspaper *Dagens Næringsliv*. Moreover, in the following week, the parliamentary members of the Minister's party publicly dismissed the quota idea.

On March 8, 2002 (event 2), however, the coalition government surprisingly proposed gender quota legislation and it promised compliance by government-controlled firms within one year. The proposal was submitted to Parliament on June 16, 2003 (event 3). It contained a sunset provision specifying that voluntary corporate compliance within two years would cause the amendment to be cancelled. As expected, the Parliament passed the proposal in November of 2003.

ASA firms began to increase female board representation following the quota proposal in 2002. This resulted in 15% female directors by the end of 2005, up from 5% in 2001. Moreover, three of five large government-controlled listed firms fully complied with the proposed quota by year-end 2002 (increasing the percent female directors to 40% from an average of 30% at year-end 2001).[3] Nevertheless, in other ASA, the change fell far short of the government's objective, and the newly elected Prime Minister announced on December 1, 2005 (event 4) that the quota would be enacted. His announcement did not, however, resolve the important question of sanctions. On December 9 (event 5), the Cabinet enacted the law, specifying forced liquidation as the ultimate penalty for noncompliance and giving existing ASA two years to comply.

In Section 4 below, we use the five events in Table 2 individually and jointly in our estimation of the market impact of quota-related news announcements. Note that the events imply that year-end 2001 is the last year in which the observed percent female directors is exogenous to the quota.

[3]We define government-controlled as the government holding at least one-third of the firm's common stock. The five government-controlled firms are Telenor ASA (telecommunications), Norsk Hydro ASA (oil and gas), Statoil ASA (oil and gas), DNB ASA (bank), and Kongsberg ASA (weapons).

2.2 Data sources and firm characteristics

We study the population of 402 OSE-listed and 724 unlisted ASA over the period 1998-2013. The firms and their characteristics are from the Norwegian national registry *Brønnøysund Register Centre* as compiled by Berner, Mjøs, and Olving (2013). As a control sample, in each sample year, we identify the 1% largest AS by revenue, of all AS with available accounting information, excluding subsidiaries not reporting consolidated accounts. This yields 3,473 unique "Large AS".[4] We allocate each sample firm to one of ten industry sectors. Listed ASA are predominantly in Offshore/Shipping, Telecom/Technology, Manufacturing, and Other Services; unlisted ASA in Financial Services, Telecom/Technology and Other Services; and Large AS in Wholesale/Retail, Construction, Manufacturing, and Other Services. Wholesale/Retail is the largest sector by number of sample firms, while Offshore/Shipping and Financial Services are the largest sectors by value.

Table 3 reports bi-annual firm and board characteristics. In an average year, there are 174 listed ASA, 255 unlisted ASA, and 987 Large AS. As shown in Column (3), measured by the average book value of total assets, Large AS are of a similar size as unlisted ASA and about one-third of listed ASA. By average revenue, however, Large AS are more than twice the size of unlisted ASA and about half the size of listed ASA (Column 1). Column (5) reports the percent female directors on the board. For ASA (both listed and unlisted), it increases from about 3% to 40% by 2008, while the increase for Large AS is from 7% to only 13% by 2013. Thus, it appears that the quota had negligible spill-over effects on the gender representation of unregulated Large AS boards.

The modest increase in the percent female directors in Large AS is similar to the general trend throughout western economies. For example, by 2013, the fraction of female directors was on average 18% in EU large publicly traded firms and 17% in US Fortune 1000 firms.[5] Columns (6) and (7) of Table 3 also list the average percent of board chairs and CEOs that are females, respectively. By 2013, the percent female chairs is 9% for listed ASA, 15% for unlisted ASA, and 5% for Large AS, up from 1%-3% in 1998. The percent female CEOs starts at about 2% in 1998 and ends in 2013 at 5% for listed ASA, 10% for unlisted ASA, and 7% for Large AS.

Table 4 summarizes the definition of variables used throughout the empirical analysis. A key such

[4] To be included, a Large AS must have positive values for total assets and revenue, and non-negative values for long-term assets, current assets, long-term and short-term debt. Moreover, we require current assets≥cash and total assets≥working capital (current assets-current debt).

[5] Source: *http://ec.europa.eu./justice/gender-equality*, and *https://www.2020wob.com.*

variable, *Shortfall*, is defined as the difference between the mandated (Table 1) and the actual fraction of female directors—a continuous variable between zero and 0.5 (measured at year-end). *High shortfall* is a dummy with a value of one if *Shortfall* is at or above its cross-sectional median. We contrast *High shortfall* firms with low-shortfall firms, where *Low shortfall = 1-High shortfall.* Moreover, we sort the ASA on their year-end 2001 board composition into two mutually exclusive groups, $Zero_{2001}$ and Pos_{2001}. The former is a dummy indicating an ASA has an all-male board in 2001, while the latter indicates a board with at least one female director. These dummies are important as 2001 is the last year in which *Shortfall* is exogenous to the quota. In 2001, 460 (83%) of 555 ASA have all-male boards.[6]

3 Shareholder actions to minimize quota-related costs

If firms and their owners perceive mandatory gender-balancing to be costly, they have an incentive to undertake actions that minimize overall gender-balancing costs. Evidence on the extent of such actions helps reveal whether firms and their shareholders viewed the quota as costly. We consider three forms of observable actions by ASA, presented in increasing order of shareholder-borne costs: (1) expand board size to make room for the new females without losing existing male directors, (2) use the director election process to maintain board CEO experience and director network power, and (3) opt out of the quota by converting the firm's legals status to AS.

3.1 Expanding board size to make room for women?

In 2001, the last year prior to the beginning of the quota-related legislative process, the average ASA board had five shareholder-elected directors. Recall from Table 1 that the quota requires a five-member board to have at least two women, while an eight-member board must have at least three women. Thus, if shareholders of a five-member board view the retention of all current male directors as important, they have the option to increase board size to eight directors. The greater board size comes with shareholder-borne costs in terms of additional director fees and possibly less efficient board decision processes (Jensen, 1993). This type of cost-increase places an upper limit on the costs of alternative strategies for minimizing shareholder-borne costs of complying with the quota.

[6]While not shown, the likelihood of an all-male board in 2001 decreases in firm value and is lower for government-controlled firms. Moreover, it is higher for firms in the Offshore/Shipping and Telecom/Technology sectors.

Figure 1 indicates that ASA board size did *not* change much over the sample period. In Panel A, the average board size remains at five directors throughout, also after the fraction of female directors reached 0.4 in early 2008. Panel B, which compares the frequency distribution of board size in 2001 and 2008, also show a relatively stable dispersion around the mean (and median) of five directors. Since board size changes are likely affected by a number of firm-specific factors, the first two columns of Table 5 present coefficient estimates in the following panel regression for firm i in year t:

$$Board\ size_{i,t} = \alpha + \gamma_1 D^{Treat} D^{Comply}_{i,t} + \gamma_2 D^{Treat}_{i,t} + \gamma_3 D^{Comply}_{i,t} + \gamma_4 \mathbf{X}_{i,t} + \epsilon_{i,t}, \tag{1}$$

where the dummy variables D^{Treat} and D^{Comply} indicate treatment (an ASA firm) and compliance (year 2008 and onwards), respectively.[7] In column (1), the control group is Large AS. In column (2), the treatment group is $Zero_{2001}$ and the control group is Pos_{2001}. Below, we use the panel estimation structure of Eq. (1) also for examining the effect of the quota on board CEO experience, board network power, and firm accounting performance.

In Eq. (1), the vector $\mathbf{X}$ of control variables, which are defined in Table 4, includes *Firm age* (log of firm age), *Total assets* (log of total assets), *ROA* (EBIT/total assets), *Leverage* (ratio of total debt to total assets), *Largest owner* (percent ownership of the largest shareholder), a dummy *Listed* indicating an OSE-listing, and industry sector dummies. The sample, which comprises 685 unique ASA and 2,627 unique Large AS, 2002-2013, excludes financial firms and AS that are registered as ASA at some point during the sample period.

In Column (1) of Table 5, the significantly positive coefficient on D^{Treat} indicates that boards in ASA are on average larger than boards in Large AS. Moreover, the coefficient on D^{Comply} is significantly negative, suggesting that the board size of Large AS declines following 2007. Most important, however, is the insignificant coefficient γ_1 on the interaction variable $D^{Treat}D^{Comply}$. This shows that any change in ASA board size after compliance is not different from that of Large AS. As to the impact of the control variables in X, board size is increasing in *Firm age* and *Total assets* and decreasing in *ROA* and *Largest owner*.

In Column (2), the sample is limited to ASA. Again, the coefficient γ_1 on the interaction variable $D^{Treat}D^{Comply}$ is statistically insignificant. That is, board size changes after the implementation of

[7] Using the actual year of compliance for ASA rather than the post-compliance period yields similar inferences throughout.

the quota are indistinguishable between firms with large and small constraints imposed by the quota. However, $\gamma_2 < 0$, showing that firms with pre-quota all-male boards tend to have fewer directors than those with gender-balanced boards.

In conclusion, there is no evidence that firms changed board size in response to the gender quota. Shareholders did not seem to find it beneficial to expand the board to make room for women without having to let go of male directors. This suggests that the perceived costs of the quota were *lower* than those of board expansion. We next examine whether shareholders were able to maintain the overall pre-quota level of board CEO experience.

3.2 Maintaining board CEO experience?

As shown in Table 3 above, the percent female CEO averages from 3% to 5% across ASA and Large AS, indicating a limited supply of female directors with CEO experience. Thus, maintaining the pre-quota overall board CEO experience *without* expanding board size requires shareholders to retain male directors with CEO experience and/or compete for females who are/were CEOs. In the analysis below, in any given sample year, director CEO experience means that the director is a current outside or past CEO of an ASA or Large AS anytime back to 1998. We exclude current inside CEOs since they are prohibited by law to sit on ASA boards after 2010. However, whether or not an inside CEO sits on the board, his/her experience will always be reflected in the board room discussion.

The dependent variable in the diff-in-diff panel estimates in columns (3) and (4) of Table 5 is the fraction of the firm's directors with *CEO experience*. Interestingly, the coefficient γ_1 on $D^{Treat}D^{Comply}$ in Column (3) is significantly negative, indicating that board CEO experience of ASA falls after compliance relative to Large AS. In Column (4), this relative drop in CEO experience is confirmed for ASA with all-male boards in 2001: $\gamma_1 < 0$ when the control group is Pos_{2001}.

Since our measure of CEO experience implies a mechanical increase in CEO experience as the look-back period increases over the sample period, D^{Comply} generates a positive and significant coefficient. For robustness, we re-estimate the regression in columns (3) and (4) with an alternative fixed three-year look-back window for CEO experience. While not shown here, these regressions generate $\gamma_{<}0$, while D^{Comply} becomes insignificant. Thus, the inference of a relative decline in CEO experience following quota compliance holds for this measure of short-run experience as well.[8]

[8]Of the control variables in X, CEO experience tends to increase with *Total assets* and *Largest owner* and decrease with

In sum, when complying, ASA constrained by the quota experienced a reduction in overall board CEO experience relative to unconstrained Large AS as well as less constrained ASA. It remains unclear why these ASA, which generally chose to maintain a five-member board size, did not simply exchange male directors *without* CEO experience for the new females. A consistent explanation is that the observed reduction in board CEO experience was perceived by shareholders to be less costly than a board size expansion or efforts to maintain overall CEO experience.

3.3 Maintaining board network power?

Director network power complements director CEO experience in terms of board effectiveness in providing valuable advise to management. Intuitively, while CEO experience measures a director's intrinsic business knowledge, the size of the director's network both amplifies this knowledge and likely enhances powers of persuasion. Thus, we are interested in whether quota compliance, by bringing in new females, significantly diminished not only average board CEO experience but also average director network power.

Figure 2 shows the frequency distribution of board seats held by ASA directors in the period prior to mandating of the quota (1998-2005) and following quota compliance (2008-2013). In Panel A, the sample is restricted to directors and directorships in listed ASA, while Panel B uses all ASA directors (in both listed and unlisted ASA) and their directorships in all ASA. As shown, directorships in ASA are highly dispersed and the dispersion of board seats is similar across male and female directors. Over the sample period, 85% of the directors sit on a single ASA board and 11% sit on two ASA boards. Thus, boards were characterized by directors with relatively low network power both before and after the quota.

The extant literature defines a director holding three or more board seats as "busy" (Ferris, Jagannathan, and Pritchard, 2003; Field, Lowry, and Mkrtchyan, 2013). With this definition, only 4% of the directors of Norwegian ASA are busy in an average sample year. In comparison, Cashman, Gillan, and Jun (2012) find that 25% of the directors of the much larger S&P 1500 companies are busy over the 1999-2008 period. Moreover, Field, Lowry, and Mkrtchyan (2013) report 45% busy directors in a sample of 1,100 US venture-capital backed initial public offerings, 1996-2008.

The fraction of busy ASA remains low throughout. However, Panel B shows that there is a slight increase (from 3% to 6%) for female directors in the post-quota period 2008-2013, offset by a slight decrease (from 6% to 3%) for male directors. Other than these small shifts, there is little evidence

Firm age and *ROA*.

that shareholders increased board seat concentration among ASA directors in order to avoid a decline in director network power.

For a more formal statistical analysis of network power changes, we estimate Eq. (1) using the board's average PageRank network power score as dependent variable. The results are shown in the last two columns of Table 5. The PageRank score uses board information for ASA and Large AS, totalling 19,206 unique directors and 96,251 directorship-years. Intuitively, with a total of N individual directors across firms in the economy, a director's network power or network "centrality" is the number of direct connections to the other $N-1$ directors. As explained by an example in Appendix A, PageRank instead uses the concept of "eigenvector centrality", which modifies the simple sum of network connections by giving greater weight to connections that are more important (emphasizing directors who themselves have important connections). Moreover, PageRank adds a small positive weight ($1/N$) to otherwise isolated directors (who receive a zero weight in the simple count).

As shown in Table 5, board network power tends to be higher for larger firms and publicly listed firms. Moreover, ASA boards generally are more powerful than boards of Large AS. Also, the coefficient for D^{Comply} is positive and significant, indicating that overall board power increased after 2007. However, the coefficient γ_1 for $D^{Treat}D^{Comply}$ is insignificant in both Column (5) and Column (6), suggesting that this increase happened across all firms. In conclusion, the quota did not seem to affect boards' overall network power.

We next consider the third and final shareholder option to minimize quota-compliance costs: opting out of the quota requirement by converting the firm's legal status from ASA to AS.

3.4 Converting from ASA to AS?

Converting from ASA to AS is done through a bylaw change at the annual general meeting. Since the legal corporate form (ASA or AS) does not affect the possibility to raise public and private debt or to be traded over the counter, a conversion has few implications for *unlisted* ASA. However, it forces an OSE-listed ASA to delist, thus imposing illiquidity costs on shareholders. On the benefit side, conversion to AS permits a firm to hold its own record of shareholders (thus avoiding Norway's central share ownership registry), and to combine the CEO and board chair positions. Moreover, it relaxes certain insider trading reporting rules. Accounting rules, such as the International Financial Reporting Standards (IFRS), are however linked to listing status and not to the legal form. Presumably, firms have optimally adjusted to these cost-benefit considerations prior to the quota.

The research question is whether imposing the quota constraint changes the prior equilibrium ASA status sufficiently to prompt conversion to AS. Because firms do not reliably disclose why they convert, the inference is necessarily indirect. To isolate potential quota-induced conversions, Table 6 shows exits from the ASA legal form and conversions to AS in the following year. Data sources for conversions are the *Brønnøysund* registry, supplemented with news and press releases. The table ends in 2007 since all firms complied with the quota by the end of 2008, i.e., the last year of conversion of relevance for our tests. Column (1) shows the annual number of non-financial ASA, 2001-2007. Financial firms are excluded because they were required to be incorporated as ASA until 2007. Column (2) records the number of firms leaving the ASA legal form in the following year due to mergers and acquisitions (M&A) or bankruptcy. In our cross-sectional analysis below, we exclude these firms since it is highly unlikely that the cost of merging with an acquirer or restructuring in bankruptcy is lower than the cost of increasing board size.

Column (3) of Table 6 lists the number of conversions for other reasons than those in Column (2) and which are of main interest for our analysis. Columns (4)-(5) report the fraction of the conversions in Column (3) undertaken by, respectively, *High shortfall* and *Low shortfall* firms and Column (6) the difference. As shown in Panel A, no listed ASA converts to AS for reasons other than M&A or bankruptcy. Apparently, whatever the perceived cost of the gender quota, it was lower than the benefits of listed ASA to remain public. In fact, the number of OSE-listed ASA increases each year from 2003 through 2007 (see Column 1).[9] On the other hand, in Panel B, Column (3) shows that on average 11% of the unlisted ASA convert to AS each year for reasons other than M&A and bankruptcy. As noted above, conversion by unlisted ASA is a relatively inexpensive if the firm has no plans to list. It is therefore interesting that *High shortfall* and *Low shortfall* firms exhibit a similar conversion frequency, suggesting that the female director shortfall is not a driver of these conversions.

To examine the potential determinants of the conversions in Column (3) of Panel B more systematically, Table 7 reports coefficient estimates for the following logit model:

$$Convert_{i,t+1} = \alpha + \gamma_1 Shortfall_{i,t} + \gamma_2 \mathbf{X}_{i,t} + \epsilon_{i,t}, \tag{2}$$

[9] This corrects an article in *The Economist*, headlined "Companies fled the [Norwegian] stock market as quotas were faced in" (November 15, 2014, p. 62).

where, in columns (1)-(4), the dependent variable $ConvertNext = 1$ in year $t = T - 1$ if an unlisted ASA switches to AS in year T and zero otherwise. Thus, $ConvertNext = 1$ only in the last year prior to conversion (year T), at which point the firm drops out. The unbalanced panel contains 821 firm-years for 148 unlisted ASA that convert and 127 that remain ASA. The vector $\mathbf{X}$ of control variables are the same as in Eq. (1) and Table 5 above with the addition of *Board size*. The odd-numbered columns use the continuous variable *Shortfall*, while the even-numbered columns use the dummy variable *High shortfall*.

The first four columns of Table 7 show that the conversion likelihood increases with *Largest owner* and decreases with *Total assets*. However, the coefficients on *Shortfall* and *High shortfall* are statistically insignificant. Thus, we find no support for the proposition that the conversion activity next year of unlisted ASA is associated with the female director shortfall implied by the gender quota. This conclusion is supported by Bøhren and Staubo (2014) in their Table 9, which uses the same definition of conversion as *ConvertNext*. Thus, both studies fail to find a significant correlation between board gender composition and the likelihood of converting in the following year.

Notwithstanding their table 9, Bøhren and Staubo (2014)'s main conclusion is that "[the conversion] response suggests that forced gender balance is costly" (abstract). They base this conclusion on an alternative regression model, where the dependent variable is backfilled after conversion. We provide a perspective on this alternative dependent variable in the last four columns of Table 7. Backfilling means that, when firm i converts from ASA to AS in year T, the dependent variable $ConvertBack$ is given a value of one in *all* previous years $t < T$. So $ConvertBack$ is always equal to one for firms that convert and zero for firms that do not convert at some point during the sample period.

We replicate the findings of Bøhren and Staubo (2014) in columns (7) and (8), where we show that the regressions for $ConvertBack$ generate significantly positive coefficients on *Shortfall* and *High shortfall*. However, this significance disappears when we add year-fixed effects in columns (5) and (6). Including year-fixed effects is necessary for the regression with backfilling to yield correct inferences. To see why, recall that *Shortfall* trends toward zero. As a result, the average *Shortfall* is higher for firms that drop out ($ConvertBack = 1$) than for firms that remain ($ConvertBack = 0$). As shown, it is this time trend in board gender composition—and not the gender composition of the individual board—that drives the significance of $Shortfall$ and $High\ Shortfall$ in columns (7) and (8).

Overall, the evidence in Section 3 fails to support the notion that shareholders of high-shortfall ASA viewed the gender quota to be sufficiently costly to have their firms increase board size or convert to

AS. We next turn to an analysis of whether stock market participants held a similar low-cost view of the quota constraint.

4 Market reaction to major quota news events

In this section, we estimate the stock market's valuation of the expected risk-adjusted cash flow changes caused by the law. The source of market values of equity is the OSE data service *Oslo Børsinformasjon*, as compiled by *Børsprosjektet* at the Norwegian School of Economics. Daily stock returns are computed using differences in the natural logarithm of daily closing prices, adjusted for splits and dividends. If a closing price is missing, it is replaced by the bid-ask midpoint (this occurs in twenty percent of the trading days).

4.1 Estimation methodology

Since the news events in Table 2 affect all firms simultaneously in calendar time, statistical inferences must account for the contemporaneous cross-correlation of stock returns. We follow Schwert (1981) and incorporate this cross-correlation by forming an equal-weighted calendar-time portfolio of all OSE-listed ASA. The inferences do not change if we use a value-weighted portfolio. Although in principle equivalent to estimating a system of seemingly unrelated regressions (SUR)—with a single OLS regression for each sample firm—the calendar-time portfolio approach is empirically more powerful as it greatly reduces the number of parameters that must be estimated. Below, we use a single-firm estimation approach only when performing cross-sectional regressions with individual firms' abnormal return estimate as dependent variable.

For each of the five events ($k = 1, ..., 5$) in Table 2, we estimate the portfolio's daily abnormal return parameter AR_k, using the following two alternative return-generating processes:

$$\text{Mean adjusted model}: \quad r_t^e = \quad \alpha + AR_k d_{k,t} + \varepsilon_t \tag{3}$$

$$\text{Market adjusted model}: \quad r_t^e = \quad \alpha + \beta W_t^e + AR_k d_{k,t} + \varepsilon_t. \tag{4}$$

The mean-adjusted model avoids estimation errors associated with a given factor-pricing model. In the market-adjusted model, W^e is the daily return on the MSCI stock market world index in excess of the daily 3-month US Treasury bill.[10] The estimation begins 252 trading days prior to event day k (day 0), ends on the event date, and excludes days of prior events, if any. To be included in the portfolio, a firm must have a minimum of 100 valid return observations (using bid-ask midpoints on non-trading days) and the firm must have an actual return observation on each day in the event window.

The dependent variable, r_t^e, is the daily stock return to OSE-listed ASA, converted to USD with the daily exchange rate and industry-adjusted. This industry adjustment subtracts, from each ASA USD return, the return on the corresponding Fama-French FF49 daily US industry portfolio return. There is one exception: for the Offshore/Shipping sector, we subtract the daily return in USD to a value-weighted portfolio of 49 *foreign* OSE-listed firms in this sector. These foreign firms, which are not subject to the quota, provide a superior industry-sector benchmark because they operate in the same geographic market as the OSE-listed ASA.

The dummy variable $d_{k,t}$ takes a value of one for each day in the event window and zero otherwise. We focus on a short, two-day event window (-1,0) as it is well known that test power increases when the abnormal return is concentrated within the event window (Kothari and Warner, 2007). With this two-day window, the event parameter AR_k is the average daily abnormal portfolio return over the two event days, and the two-day cumulative abnormal return is therefore $CAR_k = 2AR_k$. The t-statistic of CAR_k is $t = \frac{2AR_k}{\sigma_{2AR_k}} = \frac{AR_k}{\sigma_{AR_k}}$.[11]

In our application, where the exogenous quota events are uncorrelated with the market return, the conditional event-parameter estimation of AR_k in Table 8 is equivalent to a traditional two-step residual analysis (Thompson, 1985). In the traditional analysis, one first estimates the parameters in the return generating process and then uses the prediction errors to generate abnormal returns (MacKinlay, 1997).

Table 8 reports the portfolio estimates of $CAR_k = 2AR_k$ along with p-values in square brackets. The abnormal return labelled *Cumulative*, the last entry of panels A and B, is the cumulative abnormal return across all five events. *Cumulative* is estimated beginning 252 days prior to the first event (February 22, 2002) and ending on the day of the fifth event (December 9, 2005). This estimation redefines the dummy variable d_t to take a value of one for each of the five two-day event windows and thus $Cumulative = 10AR$.

[10] Our empirical results are unaffected by the inclusion of additional risk factors, such as the Fama-French factors and momentum.

[11] The ratio of the cumulative and its standard deviation is the same as the ratio of the average and its standard deviation. The estimates of AR_k and σ_{AR} are provided by the regression output.

4.2 Abnormal stock return estimates

Panels A and B of Table 8 list the two-day (-1,0) abnormal return estimates for the quota-related news events in Table 2 using the mean-adjusted model (Eq. 3) and the market-adjusted model (Eq. 4), respectively. Since the portfolio return is industry-sector adjusted, the abnormal return estimates represent a diff-in-diff estimation. The five event-portfolios of domestic OSE-listed firms in Column (1) contain 136 firms on average. In columns (2) and (3), we split the sample cross-sectionally into the two portfolios *High Shortfall* and *Low Shortfall*, respectively, measured at the year-end preceding each event date. In Column (2), the average number of firms in the portfolio is 77, while it is 53 in Column (3).[12] Column (4) estimates abnormal returns to a portfolio long in *High Shortfall* and short in *Low Shortfall* firms. These two portfolios are rebalanced each year-end with updated board data.

Importantly, none of the events in Table 8 generates statistically significant abnormal return estimates. This holds also for the long-short portfolio in Column (4), as well as for the five-event *Cumulative* abnormal return estimates. While not tabulated, results with an alternative 3-day event window (-1,1) yield identical statistical inferences, irrespective of the risk adjustment (available upon request). In sum, we conclude with a value-neutral valuation effect of the quota announcements.

Next, in Table 9, we use cross-sectional (OLS) regressions at the firm level in order examine whether the market reaction to quota news events depends on the shortfall of female directors. If the quota constraint is costly, firm i's abnormal return in response to event k, $2AR_{i,k}$ should be more negative the more binding the quota constraint, i.e. the greater *Shortfall*. The regression specification for each event k is thus:

$$2AR_{i,k} = \alpha_k + \gamma_{1,k} Shortfall_{i,k} + \gamma_{2,k}\mathbf{X_{i,k}} + u_{i,k}. \quad i = 1, \ldots, N. \tag{5}$$

Here, the vector of controls $\mathbf{X}$ includes *Largest owner* and *Total assets*, as well as three new controls (also defined in Table 4): *Codetermination* (a dummy indicating that quota-induced females and employee directors together have a majority of the board seats), *Risk* (the firm's daily stock return volatility in the year prior to the event), and a dummy indicating *Government control*. *Codetermination* and *Risk* are meant to capture, respectively, the possibility that new and relatively inexperienced female directors form coalitions with labor representatives on the board and that they may be excessively risk averse, to the detriment of shareholders. All variables are valued at the year-end prior to each event.

[12] We are unable to classify six of the firms in Column (1) due to incomplete board information.

The regressions in Table 9 uniformly fail to identify a significant effect of the shortfall dummy on the event returns for all five events. Thus, these regressions do not suggest any other interpretation of the quota-induced abnormal returns than the one given above—that the mandatory quota had a value-neutral short-term effect on OSE-listed firms.

4.3 Adjusting Ahern and Dittmar (2012) for return cross-dependence

With a sample of 94 OSE-listed ASA, Ahern and Dittmar (2012) report significantly negative average abnormal returns computed over the five-day window centered on February 22, 2002 (the date when the Minister of Industry and Trade first declares his support for a gender quota—see Table 2). Thus, our conclusion above differs from theirs. To examine why, we first replicate their abnormal return estimates. We then show that their t-test for the sample average abnormal return is overstated as it ignores the cross-dependence of returns. This cross-dependence is generally present when an event affects firms simultaneously in calendar time. As recommended by Schwert (1981), a simple way to incorporate any such cross-dependence is to form a portfolio of the sample firms and estimate abnormal returns for the portfolio, much as we do in Table 8 above. Below, we show that this portfolio estimation produces statistically insignificant abnormal returns, although similar in magnitude to those reported by Ahern and Dittmar (2012).

Panel A of Table 10 lists the original average abnormal return estimates in Ahern and Dittmar (2012) (their Table III, Panel A), computed as the average of:

$$CAR_i^{AD}(-2,2) = \sum_{\tau=-2}^{2} (r_i - r_{imatch})_\tau \,,\ i = 1, ..., 94. \tag{6}$$

Here, r_i is the return to OSE-listed ASA i on event day τ and r_{imatch} is the average return to US-listed companies in firm i's Global Industry Classification Standard (GICS) industry. Their return data are from Compustat Global for Norwegian firms and the Center for Research in Securities Prices (CRSP) for the matching firms. They report p-values only, and the standard errors for the average $CAR^{AD}(-2,2)$ shown in parentheses is inferred by us from their p-values. As listed, the sample average $CAR_i^{AD}(-2,2)$ is -2.57% with $p < 0.001$. Thus, unlike us, they conclude with a significantly negative market reaction to the February 22, 2002, quota event.

In Panel B of Table 10, we replicate this abnormal return estimation using their sample, data sources and methodology (Eq. 6).[13] The replication yields an average abnormal return estimate of -2.73% and a standard error of 0.780, which both are similar to theirs. Then, in Panel C, we re-estimate the same five-day industry-adjusted abnormal return for the 94 firms using the portfolio time-series approach in Panel A of Table 8. That is, this portfolio estimation starts 252 days prior to February 22, 2002, has the equal-weighted industry-adjusted portfolio return, $r_{p,t}^{-I} = \frac{1}{N}\sum_{i=1}^{N}(r_i - r_{imatch})_t$, as dependent variable, and uses Eq. (3) with the event dummy d_t taking a value of one over the five-day window (-2,2) and zero otherwise. However, we now use the same return data sources as Ahern and Dittmar (2012), which is Compustat Global and CRSP. This portfolio estimate is $CAR = 5AR = -2.12\%$, which is again similar to the -2.57% reported by Ahern and Dittmar (2012). However, the portfolio estimate of σ_{AR} is 0.650. As a result, the t-test of $t = 5AR/\sigma_{5AR} = AR/\sigma_{AR}$ now has a p-value of only 0.516. Thus, the average (portfolio) abnormal return estimate is statistically insignificant also when otherwise using the data and estimation procedure of Ahern and Dittmar (2012).

To reiterate, in panels A and B, the standard error of the average five-day abnormal return is computed as $\sigma_{CAR} = \frac{\sigma}{\sqrt{N}}$, where σ is the cross-sectional standard error of the $N = 94$ observations on $CAR_i^{AD}(-2,2)$ in Eq. (6), and where σ assumes cross-sectional independence of the abnormal returns. To illustrate the bias from falsely assuming cross-sectional independence, assume for simplicity that the N individual abnormal return variances σ and $N(N-1)$ pairwise return covariances ρ (between firms i and j) are cross-sectionally constant. Following Kothari and Warner (2007), we can then write

$$\sigma_{CAR} = \sqrt{\frac{1}{N}\sigma^2 + \frac{N-1}{N}\sigma^2\rho}, \tag{7}$$

and so the test bias is

$$\frac{\sigma_{CAR}}{\sigma_{CAR}\ (independence)} = \frac{\sigma_{CAR}}{\sqrt{\sigma^2/N}} = \sqrt{1+(N-1)\rho}. \tag{8}$$

In Column (1) of Table 10, the portfolio estimate of σ_{CAR} in Panel C is $5\sigma_{AR} = 3.250$. Moreover, in Panel B, $\sigma_{CAR}\ (independence) = \sigma/\sqrt{94} = 0.780$. The difference, given by the right-hand-side of Eq. (8), implies $\rho = 0.176$ on average. Interestingly, while the average pairwise cross-correlation is likely to be higher for the resource-based OSE firms than for firms on the more diversified US stock markets, this

[13] We thank Kenneth Ahern and Amy Dittmar for providing their sample of OSE-listed firms.

estimate is not far from the average pairwise correlation of stock returns of 0.110 reported by De Bodt, Eckbo, and Roll (2016) for the universe of public US manufacturing firms 1980-2011.

Finally, a caveat on the perils of using multi-day windows for politically controversial legislative events, such as the gender quota. Recall from Table 2 that, while the Minister of Industry and Trade made a highly surprising supporting statement to a tabloid news reporter on Friday February 22, 2002, he retracted his support the very next day in a high-profile interview in Norway's national business daily. While our 2-day event window (-1,0) reported in Table 8 separates these two events, the five-day event period used by Ahern and Dittmar (2012) does not. In fact, when we break down their five-day abnormal return, the average market reaction on Friday, February 22, is small and insignificant even with their unadjusted standard errors. However, the average market reaction on the subsequent Monday is negative and significant—driving their negative five-day return. This evidence suggest that, if anything, the market reacted negatively to the Saturday event that *decreased* the probability of a gender quota—the exact opposite inference of that made by Ahern and Dittmar (2012). However, this inference is also premature: our portfolio estimate of the daily abnormal return—which accounts for cross-dependence—is statistically insignificant also for Monday, February 25, 2002.

Having reconciled the prior announcement return evidence with ours, we next examine potential long-run valuation effects of the quota.

5 Gender balancing and long-run performance

The event study analysis estimates the immediate market reaction to news affecting the probability of a gender quota. This market reaction reflects the anticipated quality and impact of the new female directors that boards are forced to appoint. In this section, we examine long-run stock and accounting performance as the market observes who in fact is appointed to boards. Actual board changes that deviate substantially from market expectations may give rise to abnormal stock price changes, which we estimate using both monthly and annual data. The calendar-time stock portfolios end in 2008, while the accounting performance analysis uses data through 2013. As above, we contrast results for portfolios of ASA with high- and low shortfall of female directors throughout to capture the different constraints imposed by the quota.

5.1 Performance of long-short portfolios sorted on female director shortfall

Table 11 reports factor loadings and the abnormal return parameter α estimated using the monthly returns on two equal-weighted portfolios of OSE-listed firms consisting of ASA with $Zero_{2001}$ (all-male board) and Pos_{2001} (mixed-gender board). Firms never switch portfolios, and the monthly average number of firms in these two portfolios is 98 and 32, respectively. The return cumulation starts in February of 2002—the beginning of significant public discussion of a quota—and ends in April of 2008, when all firms are in full compliance.

The constant term α is from the following return generating process:

$$r_t^e = \alpha + \beta_1 W_t^e + \beta_2 HML_t + \beta_3 SMB_t + \varepsilon_t, \tag{9}$$

where r^e is the monthly industry-adjusted USD-denominated portfolio stock return, W_t^e is the monthly return on the MSCI world stock-market index in excess of the 3-month US treasury bill, and HML and SMB are the global value and size factors from Ken French's web site. In columns (4)-(6) we add a fourth factor, MOM, which is the global momentum factor, also from Ken French.

The abnormal performance parameter α is statistically insignificant for all portfolios, even in columns (3) and (6), where the portfolio is long in firms with high shortfall of female directors and short in firms with low shortfall. Thus, Table 11 fails to provide support for the alternative hypothesis that the quota is costly for shareholders.

5.2 Female director shortfall and Tobin's Q

We next examine whether variation in *Shortfall* affects the firm's Q from 2002 through the end of 2008 or 2009. The idea here is that investors may not fully anticipate the quality of the new female directors until they are *de facto* appointed. Thus, the total valuation effect of the quota may be deferred until the board's gender composition changes. To control for exogenous variation in female directors during the compliance period, we use a two-step instrumental variable (IV) procedure inspired by Stevenson (2010) and Ahern and Dittmar (2012).

We define Q as the ratio of book value of total assets minus book value of equity plus market value of equity to book value of total assets. The market value of equity is the stock price times the number of shares outstanding (shares issued - treasury shares), using the end-of-year closing price. If a firm has more than one share class, the market value of equity is the combined market value of all share classes. We drop Q-values that are less than or equal to zero, and then winsorize the resulting observations at 1% and 99% each year. The data used to compute Q is from *Børsprosjektet*.

Before proceeding with the IV test, we estimate the following reduced-form regression, with an unbalanced panel of 239 listed ASA over the period 2002-2008:

$$Q_{i,t} = \alpha + \beta Shortfall_{i,t} + \theta_i + \tau_t + \epsilon_{i,t}, \tag{10}$$

where θ and τ are firm- and year- fixed effects, respectively. This estimation yields a statistically insignificant coefficient estimate of $\beta = 0.030$, suggesting that the actual appointment of female directors is unrelated to Q (the regression results are available upon request).

Next, in Table 12, much as in Ahern and Dittmar (2012), we identify exogenous variation in $Shortfall_{i,t}$ over the period 2002 through 2008. The purpose is to remove endogenous time-series variation in $Shortfall_t$ that arises because firms self-select the timing with which they comply with the quota (up until 2008). In the first stage of the IV estimation, we regress $Shortfall_{i,t}$ on $Shortfall_{i,T_0}$ interacted with year dummies as follows:

$$Shortfall_{i,t} = \alpha + \sum_{\tau=T_1}^{T_2-1} \beta_\tau D_\tau Shortfall_{i,T_0} + \theta_i + \tau_t + u_{i,t}. \tag{11}$$

D_τ is a year dummy and T_1 and T_2 are the beginning and end, respectively, of the estimation period. T_0 is the base-year in which the firm's board gender composition is independent of the quota and in which the exogenous $Shortfall_{i,T_0}$ is measured. This first-stage regression is shown in Panel B of Table 12. In the second-stage OLS regression shown in Panel A, we estimate the impact on $Q_{i,t}$ of the predicted shortfall $\widehat{Shortfall}_{i,t}$ from the first-stage, including firm- and year-fixed effects:

$$Q_{i,t} = \alpha + \beta \widehat{Shortfall}_{i,t} + \theta_i + \tau_t + \epsilon_{i,t}, \qquad t = T_1, .., T_2. \tag{12}$$

Column (1) of Table 12 shows our main regression results. Recall that 2001 is the last year in which the cross-sectional distribution of female directors is unaffected by the subsequent quota legislative events (which began in 2002). Thus, $T_0 = 2001$, $T_1 = 2002$ and $T_2 = 2008$ and the regression ends in 2008—the first year in which all firms are in compliance. In Panel A of Column (1), the slope coefficient estimate for $\widehat{Shortfall}$ is a statistically insignificant 0.750. As shown in Column (2), this conclusion is robust to moving the exogenous *Shortfall* one year back, to $T_0 = 2000$. While not shown, the conclusion is unchanged also if we include firm-specific control variables, such as the natural logarithm of book value of assets, book leverage, and board size, in the second stage. In sum, these estimates imply that the development in Q from 2002 through 2008 is unrelated to the instrumented female director shortfall.

5.3 Ahern and Dittmar (2012) and Tobin's Q

Ahern and Dittmar (2012) report a negative and significant effect of the fraction of female directors on industry-adjusted Q when using a two-stage IV test much like the one in Table 12. Most important, they select year-end 2002 to measure the exogenous board-composition ($T_0 = 2002$) in their first-stage estimation, which is otherwise similar to Panel B in Table 12. In columns (3)-(5), we replicate their estimation, which starts in $T_1 = 2003$ and ends in $T_2 = 2009$, thus giving investors one more year to realize the valuation impact of the forced board changes. Moreover, in Column (3), we also apply their base-year of $T_0 = 2002$. This results in a coefficient estimate for $\widehat{Shortfall}$ of 1.91 in Panel A and which is almost identical to the point estimate of -1.92 reported by Ahern and Dittmar (2012) and with a similar significance level.[14]

In columns (4) and (5), we further show that the significance of the slope estimate in Column (3) is driven by the choice of $T_0 = 2002$ and not by the alternative estimation period of 2003-2009. Specifically, in Column (4), we apply $T_0 = 2001$ to this estimation period, and in Column (5) we use $T_0 = 2000$. The estimate of the coefficient β is again statistically insignificant in both these two columns. Thus, the only model specification that produces a significant coefficient estimate for $\widehat{Shortfall}$ is the one in Column (3), where $T_0 = 2002$ in the first stage instrumentation.

Figure 3 illustrates the dramatic effect of shifting the base-year from $T_0 = 2001$ in Column (4) to $T_0 = 2002$ in Column (3) of Table 12. Each line in the figure is a firm-level pathway from 2003 through 2009 for the annual fitted value of $\widehat{Shortfall}$. In Panel A of Figure 3, this fitted value is based on $T_0 = 2001$ (corresponding to Column 4 in Table 12), while it is based on $T_0 = 2002$ in Panel B (corresponding to Column 3 in Table 12). Each pathway moves the initial and presumably exogenous base-year female director shortfall towards full quota compliance (zero shortfall) at the speed of the sample-wide average

[14] The sign of our coefficient estimate is switched because we use the shortfall of female directors, while they use the (near-inverse) percent female directors. Also, Ahern and Dittmar (2012) use industry-adjusted Q, which we do not (Gormley and Matsa, 2014). Our results are, however, unchanged when using industry-adjusted Q.

firm-year shortfall change. The idea behind the instrumentation is that these lines are unaffected by firm-specific shortfall deviations, which are possibly strategic and therefore endogenous.[15]

As illustrated, the cross-sectional dispersion of the pathways increases noticeably from Panel A to Panel B of Figure 3. The reason for this increased spread is the budding compliance by ASA (and, in particular, government-controlled ASA) in response to the initial quota-related events in 2002, and which invalidates Ahern and Dittmar (2012)'s choice of $T_0 = 2002$. As Table 12 demonstrates, with an appropriate base-year for the instrumentation ($T_0 < 2002$), there is no evidence of an impact of *Shortfall* on Q in the second stage.

Finally, we turn to an estimation of the impact of the quota on operating profitability.

5.4 Gender balancing and operating profitability

In this section, as a complement to the above market-value based analysis, we examine the effect of the quota on firms' operating performance by estimating the following diff-in-diff panel regression:

$$ROA_{i,t} = \alpha + \gamma_1 D^{Treat} D^{Comply}_{i,t} + \gamma_2 D^{Treat}_{i,t} + \gamma_3 D^{Comply}_{i,t} + \gamma_4 \mathbf{X}_{i,t} + \epsilon_{i,t}, \tag{13}$$

where we add *Listed* to the vector of control variables $\mathbf{X}$ in Table 6. The coefficient estimates are reported in Table 13. As shown in Column (1), operating performance is on average lower for ASA than Large AS. Moreover, ROA increases with firm age and size, and decreases with board size and leverage. As before, the key variable of interest is $D^{Treat} D^{Comply}$. This variable receives a statistically insignificant coefficient estimate in both Column (1) and Column (2). Thus, there is no evidence that quota compliance affected operating performance for ASA relative to Large AS, or for ASA with pre-quota all-male boards relative to those with already mixed-gender boards.

Matsa and Miller (2013) find that the quota led to a relative increase in labor costs and decrease in operating profitability. Their estimation relies on a post-compliance period that starts in 2007. We instead use 2008 as the year of full quota compliance since this is the first year in which all regulated firms actually complied. More importantly, their sample period is limited to 2003-2009. The effect of the shorter sample period is shown in Column (3) of Table 13, which uses the regression specification in

[15] While the number of unique ASA is 227 in the figure, the number of distinct pathways is much smaller (13 in Panel A and 12 in Panel B of Figure 3) simply because firms with the same initial shortfall must have the same estimated pathway until 2009.

Column (1) but limits the sample to 2003-2009.

As in Matsa and Miller (2013), the coefficient on $D^{Treat}D^{Comply}$ in Column (3) is now significantly negative, suggesting that ASA had lower ROA in 2008-2009 than in 2003-2007 compared to Large AS. However, this inference does not hold when the sample is extended one year forward to 2010 (Column 4). Also, using the actual year of compliance in columns (5) and (6) yields the same inference of an insignificant coefficient for $D^{Treat}D^{Comply}$ for the 2003-2009 sample period (as well as the full sample period, not shown here). Thus, the conclusion of Matsa and Miller (2013) is not robust to an extension of their sample period, or to using the actual compliance year in the analysis. Overall, we conclude that the quota had no statistically significant impact on firms' operating profitability.

In sum, the analysis of short- and long-run performance in this section suggests that the gender quota did not significantly affect the market value or the operating profitability of the typical ASA. This finding is consistent with the lack of corporate actions that could reduce compliance costs, as documented earlier in Section 3.

6 Conclusions

Norways pioneering mandatory gender quota initiated a wave of subsequent legislation throughout Europe instituting gender-balancing of the boards of public limited liability companies. This paper presents a battery of tests on the economic consequences of Norway's quota, which in economic terms pitched the value of grater director independence against the loss of director experience. The tests range from short- and long-term market valuation estimates to corporate actions taken to minimize shareholder-borne costs of the quota (increase board size, retain highly qualified males, change corporate legal form). Our diff-in-diff approach contrasts treated firms (ASA) with a control sample of large AS, as well as treated firms facing a hard quota constraint (all-male board) versus a weaker constraint (some female directors) prior to quota implementation.

A unique methodological aspect of this paper is to link our market valuation estimates to evidence on shareholder actions to reduce or avoid perceived costs of forced gender-balancing. Surely, if investors and firms anticipated large negative effects—as prior research suggests—one would expect to see evidence of corporate evasive actions. The perhaps simplest such action is to increase board size in order to fill the quota without having to let go of valued male directors. Another action is to convert legal form from

ASA to AS, which is costly for OSE-listed ASA (as it forces delisting) but not particularly costly for unlisted ASA with no plans to list. In this argument, the shareholder-borne costs of increasing board size without losing male directors (which would entail going from five to eight members for the typical ASA), or converting to AS, place an upper bound on the costs of the quota and therefore also on the market capitalization of those costs.

In the first part of the paper, our diff-in-diff tests show that regulated firms did *not* increase board size relative to large AS or ASA with a lower quota constraint. Second, none of the listed ASA converted to AS. Third, while unlisted ASA regularly switch legal form (about ten percent go from ASA to AS annually throughout our sample period), we show that this conversion activity is statistically unrelated to the quota constraint (the quota-induced shortfall of female directors) when accounting for the general time-trend in board gender composition. This evidence does not leave much room for arguing that firms perceived the quota to be particularly costly. We do show, however, that although board network power does not decline, board CEO experience does decline post-compliance. Nevertheless, it appears that the cost of this decline in experience was insufficient to prompt firms to either increase board size or convert to AS.

In the second part of the paper, we estimate effects of the quota on the market values of listed ASA, and we examine the operating performance of ASA versus large AS. We again single out firms with all-male boards. Our market value analysis fails to reject the null of a value-neutral effect of the quota constraint—even for firms with all-male boards. While this evidence contrasts with prior evidence of negative valuation effects, we also show that the prior evidence is consistent with a zero valuation effect after two econometric adjustments. The first corrects for cross-dependence in stock returns that arises because the event affects all firms simultaneously in calendar time. The second adjustment creates an instrument for the time-series variation in the percent female directors that is truly exogenous to quota implementation. Finally, our operating performance analysis fails to find an impact of the quota on the rate of return on assets. While earlier research documents negative operating performance, this prior evidence hinges on a short sample period post-quota-implementation ending in 2009.

On a fundamental level, our empirical analysis suggests that the supply of professionally qualified female directors was sufficiently large in Norway to justify a neutral corporate and market reaction to the quota. Thus, our answer to the question posed in the title of this paper is that neither investors nor the firms themselves appear to have viewed forced gender balancing as particularly costly.

A PageRank computation of network power

Consider the network of $N = 8$ directors in four firms illustrated in Appendix Figure 1. Each node in the figure is a director, and two nodes are connected with a line if the two directors sit on the same board. Thus, C and D are directors of one firm, while D and E sit on the board of another. A third firm has A, B and H as directors, while the directors of a fourth firm are F, G and H. Appendix Table 1 reports the eigenvector centrality and PageRank scores for each node in this network.

Let $\mathbf{A}$ denote this network's $N \times N$ adjacency matrix, where entry a_{ij} is equal to 1 if node i is connected to node j and zero otherwise. Thus, $\mathbf{A}$ is as follows:

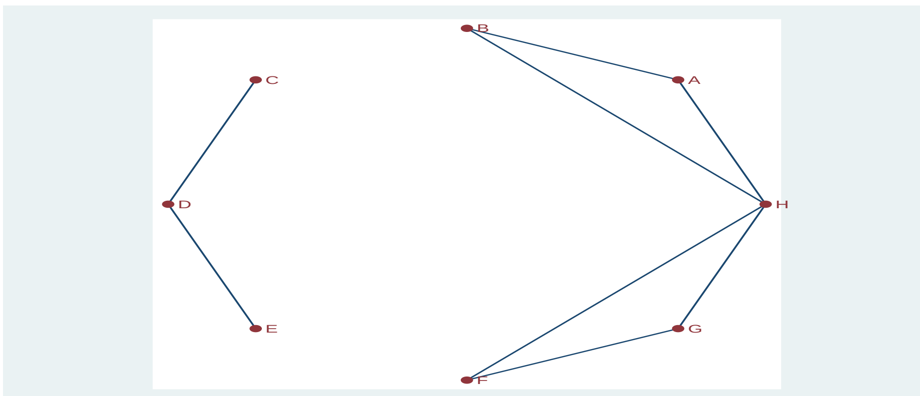
	A	B	C	D	E	F	G	H
A	0	1	0	0	0	0	0	1
B	1	0	0	0	0	0	0	1
C	0	0	0	1	0	0	0	0
D	0	0	1	0	1	0	0	0
E	0	0	0	1	0	0	0	0
F	0	0	0	0	0	0	1	1
G	0	0	0	0	0	1	0	1
H	1	1	0	0	0	1	1	0

$\mathbf{A}$'s eigenvector centrality is the eigenvector corresponding to the largest eigenvalue of $\mathbf{A}$, which is 2.561 in this example. To compute the PageRank, first normalize each column in $\mathbf{A}$ by the column sum, which yields the following normalized adjacency matrix $\mathbf{A}^*$ (where the elements of each column sum to one):

	A	B	C	D	E	F	G	H
A	0	1/2	0	0	0	0	0	1/4
B	1/2	0	0	0	0	0	0	1/4
C	0	0	0	1/2	0	0	0	0
D	0	0	1	0	1	0	0	0
E	0	0	0	1/2	0	0	0	0
F	0	0	0	0	0	0	1/2	1/4
G	0	0	0	0	0	1/2	0	1/4
H	1/2	1/2	0	0	0	1/2	1/2	0

Second, construct the 8×8 transition matrix $\mathbf{B}$, where each entry is $b_{ij} = 1/8$, and combine $\mathbf{A}^*$ and $\mathbf{B}$ to produce the matrix $\mathbf{C} = [1-\delta]\mathbf{A}^* + \delta\mathbf{B}$. Here a typical scaling factor is $\delta = 0.15$. Matrix $\mathbf{C}$ is positive and its largest eigenvalue is 1 (by the Perron-Frobenius Theorem). The vector of PageRank scores for each of the eight directors is the eigenvector in $\mathbf{C}$ with an eigenvalue of one, normalized with the sum of the elements in the vector.

Appendix Figure 1
Network example with eight directors in four firms



Appendix Table 1
Eigenvector and PageRank values for the network example

Node	Eigenvector	PageRank
A	0.3941	0.1064
B	0.3941	0.1064
C	0.0000	0.0963
D	0.0000	0.1824
E	0.0000	0.0963
F	0.3941	0.1064
G	0.3941	0.1064
H	0.6154	0.1996

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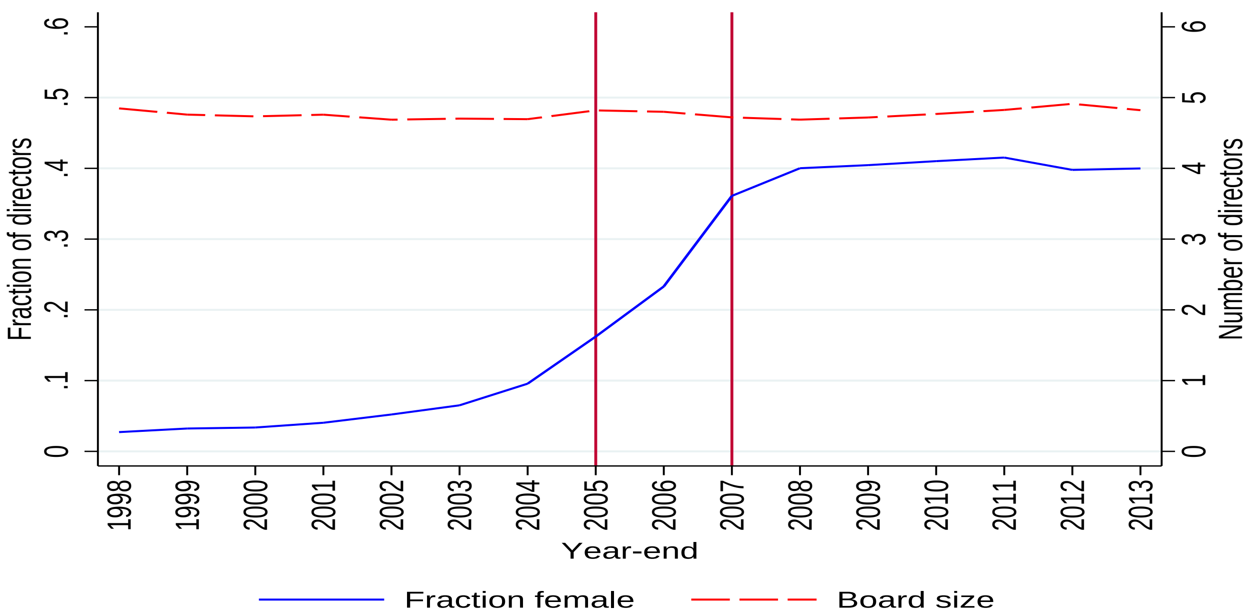
Figure 1
ASA board size and fraction of female directors

Panel A shows the average board size (number of shareholder-elected directors) and fraction of female directors for the population of 1,126 Norwegian ASA, 1998-2013. Panel B plots the board size frequency distributions in 2001 (N=555) and 2008 (N=395). The two vertical lines mark year-end 2005, when the quota was signed into law, and year-end 2007, when the formal quota compliance period ended. All ASA firms complied by year-end 2008.

A: Average ASA board size and fraction of female directors



B: ASA board size frequency distribution in 2001 and 2008

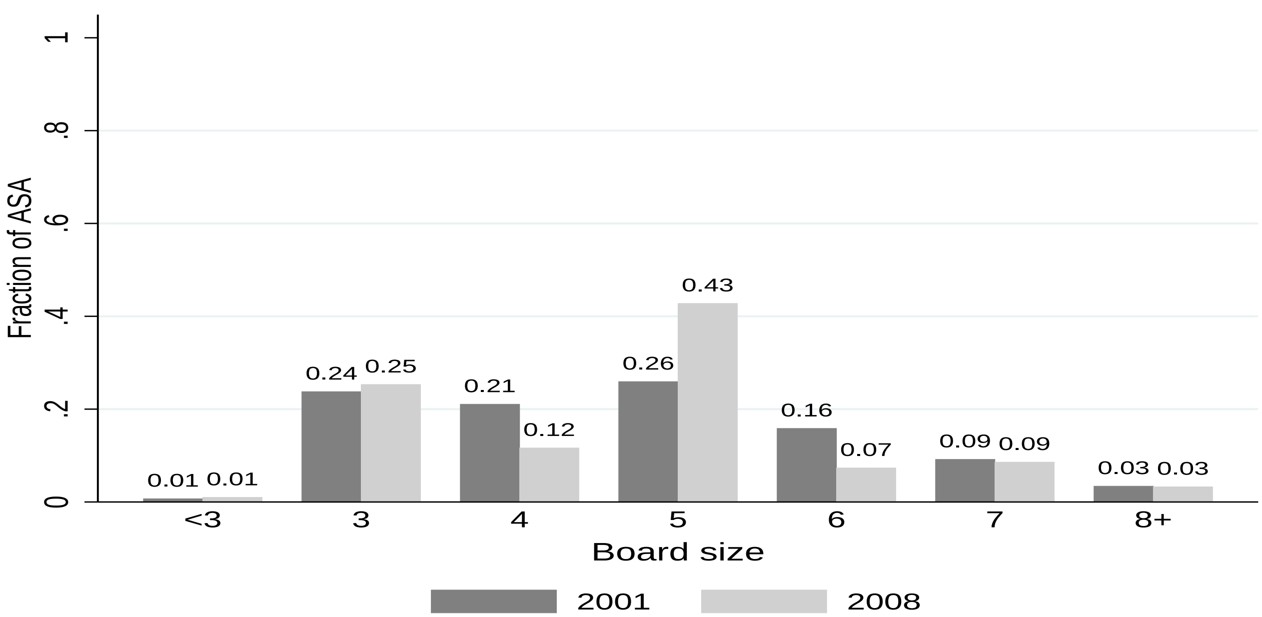
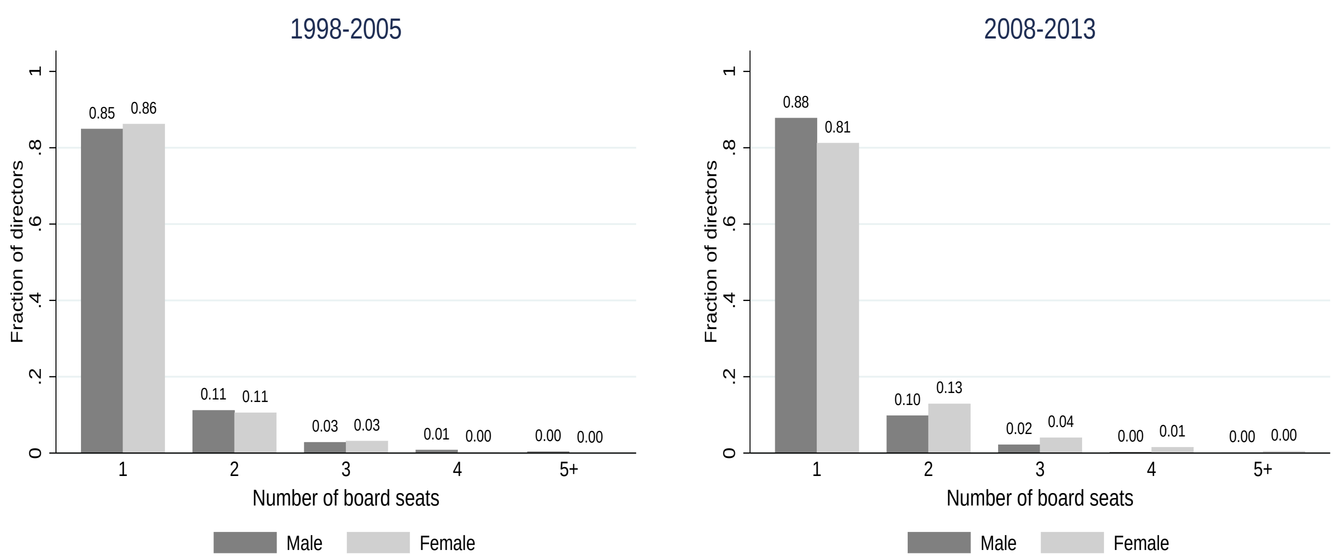


Figure 2
Frequency distribution of the number of board seats held by ASA directors, 1998-2013

The figure plots frequency distributions of the number of board seats held by ASA directors. Panel A counts board seats in listed ASA, while Panel B counts board seats in all ASA (listed and unlisted). Five and more board seats are reported as 5+. Each panel shows the pooled distribution for the periods 1998-2005 (before the quota is mandated formally) and 2008-2013 (after full compliance). The sample is 402 listed ASA and 867 unlisted ASA, 1998-2013.

A: Number of board seats in listed ASA by listed ASA directors



B: Number of board seats in all ASA by all ASA directors

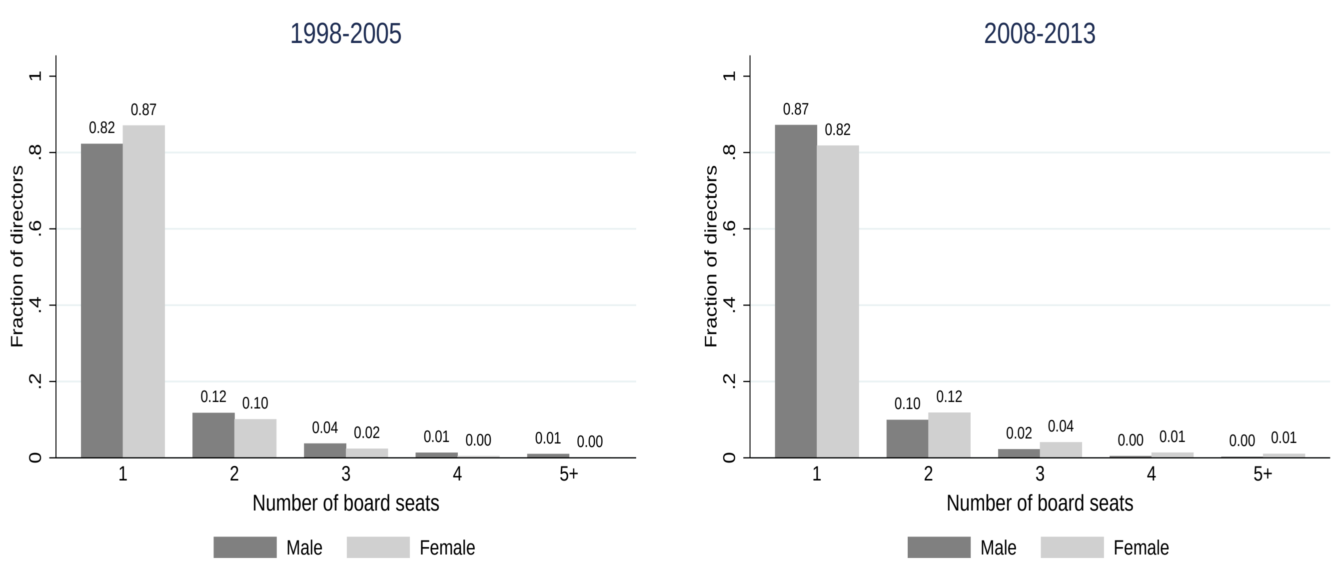


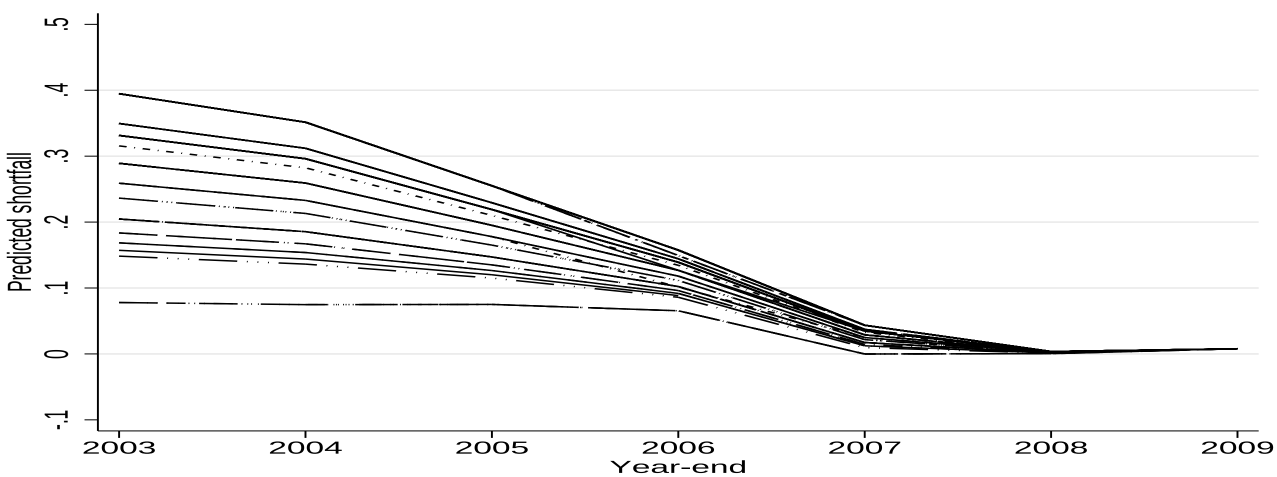
Figure 3
First-stage instrumentation of female director shortfall

The figure shows the predicted shortfall female directors relative to the quota requirement, $\widehat{Shortfall}_{i,t}$, estimated from the following first-stage regression:

$$Shortfall_{i,t} = \alpha + \sum_{\tau=2003}^{2008} \beta_\tau D_\tau Shortfall_{i,T_0} + \theta_i + \tau_t + u_{i,t},$$

where D_τ is a year dummy, and θ_i and τ_t are firm and year fixed effects, respectively. $Shortfall_{i,T_0}$, which is the female director shortfall exogenous to the quota, is measured at year-end $T_0 = 2001$ in Panel A and $T_0 = 2002$ in Panel B. Each estimated line moves the initial $Shortfall_{i,T_0}$ towards zero (full compliance) at the speed of the sample-wide average change. The sample is 227 listed ASA, 2003-2009.

A: Predicted shortfall $\widehat{Shortfall}_{i,t}$ with $Shortfall_{i,T_0}$ measured at year-end $T_0 = 2001$



B: Predicted shortfall $\widehat{Shortfall}_{i,t}$ with $Shortfall_{i,T_0}$ measured at year-end $T_0 = 2002$

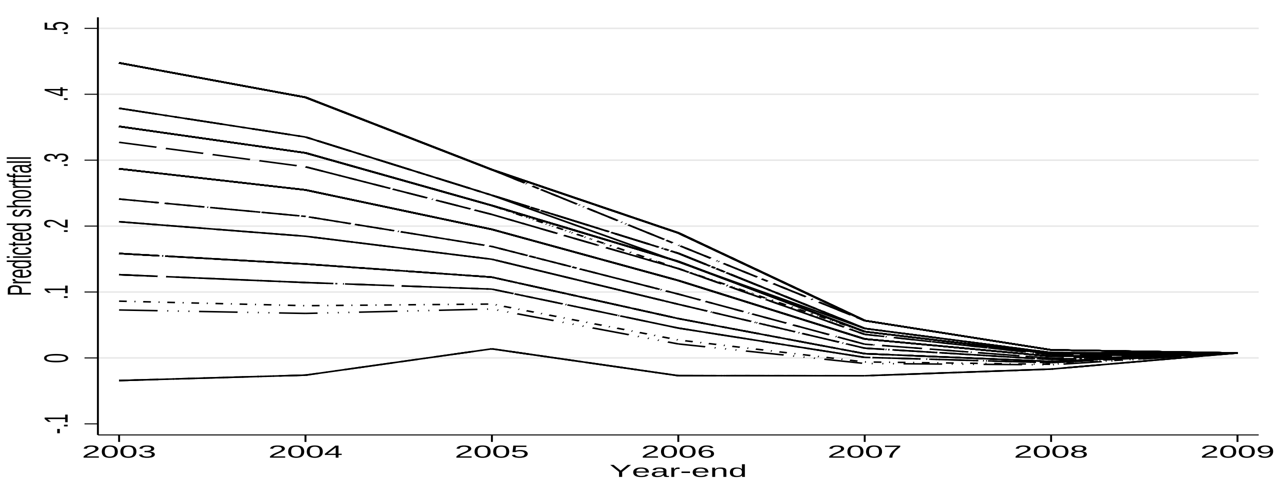


Table 1
Female directors required by Norway's board gender quota

The table shows how the required number and fraction of female directors varies with board size, defined as the number of shareholder-elected directors.

Board size	Required number of female directors	Required fraction of female directors
3	1	0.33
4	2	0.50
5	2	0.40
6	3	0.50
7	3	0.43
8	3	0.38
9	4	0.44
10	4	0.40
>10	>4	≥0.40

Table 2
Major news events increasing the probability of a mandatory board gender quota.

A: Events leading to approval by Parliament of the corporate board gender quota

(1) February 22, 2002: The Minister of Trade and Industry surprisingly supports a gender quota in a newspaper interview (*Verdens Gang*).

The next day, on Saturday February 23, 2002, the same Minister publicly retracts his support in a newspaper interview (*Dagens Næringsliv*). One week later, the parliamentary members of his party rejects a quota.

(2) March 8, 2002: The Cabinet surprisingly proposes a board gender quota to be signed into law in 2005. The proposal contains a sunset provision, cancelling the amendment if firms comply voluntarily by 2005. The Cabinet promises compliance by government-owned firms within one year (*Dagens Næringsliv*).

(3) June 16, 2003: Parliament receives the gender quota proposal from the Cabinet (*Aftenposten*).

The quota proposal is approved by Parliament in November of 2003, as expected from the public debate during the fall.

B: Signing of the board gender quota into corporate law

(4) December 1, 2005: The newly elected Prime Minister declares that the Cabinet will sign the gender quota amendment into law. There are speculations that the sanctions will be a monetary fine. (*Verdens Gang*).

(5) December 9, 2005: The new Cabinet signs the quota amendment into law. It announces that the ultimate sanction for noncompliance is forced liquidation—as for any breach of Corporate Law. Existing ASA are given two years to comply (*Dagens Næringsliv*).

Table 3
Firm and board characteristics for ASA and Large AS, 1998-2013

Bi-annual firm and board characteristics for the sample of listed ASA (Panel A), unlisted ASA (Panel B), and Large AS (Panel C), 1998-2013. Columns (1)-(4) show the mean and median revenue and book value of total assets, reported in million 2013 USD and winsorized at the 1% tails. Columns (5), (6), and (7) list the average percent female directors, chairs, and CEOs, respectively. Large AS is the top 1% AS by revenue. The last row in each panel lists the pooled average across all firm-years, with the exception of column (7), which lists the average annual number of firms across the sample period.

	Revenue		Total asset		Percent female			
Year	Mean	Median	Mean	Median	Directors	Chair	CEOs	N
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	
A: Listed ASA								
1998	409	96	756	144	3.3	1.6	2.7	196
2000	347	59	727	125	4.6	1.1	1.7	193
2002	451	80	865	121	6.8	1.3	3.3	160
2004	506	74	863	117	14.4	1.9	1.9	155
2006	497	75	1,001	189	28.3	1.7	2.3	175
2008	556	151	1,279	375	40.3	3.1	2.6	193
2010	577	119	1,229	275	41.4	5.8	2.9	174
2012	665	146	1,303	361	40.7	10.1	3.1	159
2013	699	163	1,373	266	41.6	8.8	4.7	150
Average	501	89	1,008	190	24.3	3.6	2.8	174
B: Unlisted ASA								
1998	46	3	133	6	2.3	0.4	1.7	247
2000	49	3	107	7	2.8	0.8	2.2	387
2002	91	3	226	7	4.6	1.6	4.4	390
2004	99	5	256	11	7.3	3.0	3.7	334
2006	84	5	338	13	20.3	3.5	4.9	289
2008	117	4	511	21	39.8	10.4	7.0	202
2010	133	8	555	25	40.6	10.3	11.2	155
2012	241	14	778	50	38.2	8.8	10.1	91
2013	229	17	787	87	37.1	15.1	9.6	86
Average	90	4	290	11	15.2	4.2	5.3	255
C: Large AS								
1998	174	78	180	48	7.4	3.3	2.3	918
2000	132	56	169	38	8.2	2.5	3.4	943
2002	150	60	197	42	9.3	3.4	3.0	963
2004	173	69	219	46	10.8	4.5	4.3	1,003
2006	216	90	280	67	13.0	4.2	5.0	975
2008	340	143	497	116	13.1	4.1	5.2	1,019
2010	275	114	431	99	14.0	5.7	6.7	1,000
2012	274	116	442	103	14.0	5.9	7.0	1,101
2013	293	118	479	103	13.2	5.5	7.2	1,158
Average	224	90	319	69	11.6	4.4	4.9	987

Table 4
Summary of firm and board characteristics used in the empirical analysis

Name	Definition (data sources: Brønnøysund Register Centre and, for Q, Børsprosjektet)
A: Variables central to the difference-in-difference tests	
D^{Treat}	Dummy indicating the treatment group, either comprising all ASA (with Large AS as control group) or Zero_{2001} (with Pos_{2001} as control group).
D^{Comply}	Dummy indicating years 2008-2013, when all ASA fully complied with the quota.
ASA	Public limited liability company ("Allmenaksjeselskap"), regulated by the quota.
Large AS	The 1% largest limited liability companies ("Aksjeselskap") by revenue. Not regulated by the quota.
Zero_{2001}	An ASA with zero shareholder-appointed female directors in 2001.
Pos_{2001}	An ASA with at least one shareholder-appointed female director in 2001, Pos_{2001}=1-$Zero_{2001}$.
B: Board characteristics	
Board size	The number of shareholder-appointed directors on the board.
CEO experience	Dummy indicating that a director is a current outside CEO or a past CEO of an ASA or Large AS going back to 1998.
Shortfall	The difference between the fraction of female directors required by the quota (see Table 1) and the actual fraction of female directors on the board.
High shortfall	An ASA with $Shortfall$ at or above the median. In 2007, the median $Shortfall$ is zero and we require $Shortfall > 0$.
Low shortfall	An ASA with $Shortfall$ below the median, *Low shortfall* = 1-*High shortfall*.
Network power	Director PageRank score. See Appendix A for the computation of this eigenvalue-based network score.
C: Firm characteristics	
Firm age	Natural logarithm of the firm's age since incorporation.
Total assets	Natural logarithm of the book value of total assets.
ROA	Ratio of earnings before interest and taxes (EBIT) to total assets.
Leverage	The ratio of book value of total debt to total assets.
Largest owner	Percent ownership of the largest shareholder.
Listed	Indicates that an ASA is listed on the Oslo Stock Exchange (OSE).
Government control	A dummy indicating that the government owns 30% or more of a listed ASA.
Codetermination	A dummy indicating that employee representatives and the female directors required by the quota have a majority of the board seats.
Risk	The firm's daily stock return volatility in the year prior to the event.
Q	Ratio of total assets minus book value of equity plus market value of equity to total assets.
Industry dummies	Indicates the firm's industry sector. There are ten different industry sectors.

Table 5
Estimates of post-quota changes in board size, CEO experience, and network power

The table reports coefficient estimates from the following panel OLS regression for firm i in year t:

$$Y_{i,t} = \alpha + \gamma_1 D_{i,t}^{Treat} D_{i,t}^{Comply} + \gamma_2 D_{i,t}^{Treat} + \gamma_3 D_{i,t}^{Comply} + \gamma_4 \mathbf{X}_{i,t} + \epsilon_{i,t}.$$

The dependent variable Y equals *Board Size* in columns (1) and (2), the fraction of the firm's directors who has CEO experience in columns (3) and (4), and the board's average director network power in columns (5) and (6). In the odd-numbered columns, the treatment group (D^{Treat}) is ASA, while the control group (1-D^{Treat}) is Large AS. In the even-numbered columns, D^{Treat} is ASA with pre-quota all-male boards ($Zero_{2001}$), while the control group is ASA with mixed-gender boards (Pos_{2001}). D^{Comply} is a dummy indicating the period 2008-2013, in which all ASA comply with the quota. The vector $\mathbf{X}_{i,t}$ is a vector of control variables. All variables are defined in Table 4. The sample, which comprises 685 ASA and 2,627 Large AS, 2002-2013, excludes financial firms and Large AS registered as ASA at some point during the sample period. Standard errors clustered by firm are reported in parenthesis. Stars indicate significance levels: *** 1%, ** 5%, and * 10%.

Dependent variable (Y):	Board size		Board CEO experience		Board power	
Treatment group (D^{Treat}):	ASA	$Zero_{2001}$	ASA	$Zero_{2001}$	ASA	$Zero_{2001}$
Control group (1-D^{Treat}):	Large AS	Pos_{2001}	Large AS	Pos_{2001}	Large AS	Pos_{2001}
	(1)	(2)	(3)	(4)	(5)	(6)
$D^{Treat} D^{Comply}$	-0.092	0.151	-0.043***	-0.101**	-0.005	-0.007
	(0.079)	(0.236)	(0.014)	(0.043)	(0.004)	(0.013)
D^{Treat}	0.355***	-0.515***	-0.012	0.052**	0.018***	0.005
	(0.095)	(0.187)	(0.015)	(0.024)	(0.003)	(0.007)
D^{Comply}	-0.161***	-0.379*	0.042***	0.105***	0.034***	0.041***
	(0.044)	(0.228)	(0.007)	(0.039)	(0.001)	(0.011)
Firm age	0.113***	0.220***	-0.008*	-0.025**	-0.002**	0.001
	(0.028)	(0.061)	(0.004)	(0.011)	(0.001)	(0.003)
Total assets	0.239***	0.233***	0.012***	0.015***	0.010***	0.006***
	(0.021)	(0.037)	(0.003)	(0.005)	(0.001)	(0.002)
ROA	-0.429***	-0.332**	-0.041**	0.019	-0.008**	-0.003
	(0.107)	(0.147)	(0.017)	(0.024)	(0.004)	(0.006)
Leverage	-0.116	0.133	0.006	0.026	-0.002	0.005
	(0.110)	(0.177)	(0.019)	(0.028)	(0.004)	(0.007)
Largest owner	-1.310***	-1.188***	0.095***	0.062*	-0.005*	-0.008
	(0.097)	(0.188)	(0.016)	(0.036)	(0.003)	(0.009)
Listed	0.028	-0.022	0.007	0.015	0.027***	0.031***
	(0.107)	(0.153)	(0.017)	(0.023)	(0.004)	(0.007)
Constant	2.503***	2.402***	0.192***	-0.022	0.089***	0.187***
	(0.313)	(0.603)	(0.049)	(0.079)	(0.010)	(0.039)
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.209	0.295	0.063	0.062	0.236	0.281
Firm-years	12318	1919	12317	1919	12030	1919

Table 6
Exits from ASA and conversions to AS

Column (1) lists the annual number of non-financial ASA, 2001-2007. Columns (2) and (3) list the number of firms in year t that exit the ASA legal form in year $t+1$. In Column (2), the reason for exit is M&A or bankruptcy. Column (3) lists the number of firms converting from ASA to AS for all other reasons than those in Column (2). Columns (4)-(6) report the conversions in Column (3) as a percent of all firms, and split by firms with high- and low shortfall of female directors. The last row in Panel B shows the average pooled across all years. The sample consists of 277 unique listed non-financial ASA (Panel A) and 456 unique unlisted non-financial ASA (Panel B). The significance for the difference in Column (7) is denoted *** 1%, ** 5%, and * 10%, and is from a two-sample t-test assuming unequal variances.

			Conversion from ASA to AS in $t+1$ for all other (unspecified) reasons				
Year t	Number of ASA at year-end t (1)	Exit from ASA due to M&A or bankruptcy (2)	Number of other conversions (3)	Percent of all firms (4)	Percent of high shortfall firms (5)	Percent of low shortfall shortfall (6)	Difference high-low shortfall (7)
A: Listed ASA							
2001	157	10	0				
2002	148	12	0				
2003	140	11	0				
2004	144	8	0				
2005	161	16	0				
2006	166	19	0				
2007	195	18	0				
B: Unlisted ASA							
2001	288	25	27	9.4	8.2	12.0	-3.8
2002	258	35	19	7.4	4.0	10.7	-6.7**
2003	221	14	19	8.6	8.0	10.8	-2.7
2004	210	45	19	9.0	6.8	16.7	-9.9*
2005	161	15	25	15.5	18.9	11.3	7.6
2006	153	27	27	17.6	27.3	7.9	19.4**
2007	126	15	12	9.5	12.9	8.4	4.5
Average				10.7	10.6	10.8	-0.2

Table 7
Determinants of the conversion likelihood for unlisted ASA

The table reports the coefficient estimates from the following logit regression for firm i in year t:

$$Convert_{i,t} = \alpha + \gamma_1 Shortfall_{i,t} + \gamma_2 \mathbf{X}_{i,t} + \epsilon_{i,t}.$$

In the first four columns, the dependent variable is $ConvertNext$, which takes the value of one in year $t = T-1$ if the firm converts to AS in year T, and zero otherwise. In the last four columns, the dependent variable is $ConvertBack$, which takes the value of one for all $t < T$ (back-filling) if a firm converts to AS in year T. Converting firms drop out of the sample in year T. The explanatory variables include *Shortfall* (odd-numbered columns) and $High\ Shortfall$ (even-numbered columns) as well as the control variables in $\mathbf{X}_{i,t}$. All variables are defined in Table 4. The sample comprises 261 unlisted non-financial ASA, 2001-2007, of which 144 convert to AS in year 2002-2008 and 127 do not convert. We exclude firms that exit the ASA legal form due to M&A and bankruptcy (listed in column (2) of Table 6). Standard errors clustered by firm are reported in parenthesis. Significance levels: *** 1%, ** 5%, and * 10%.

Dependent variable:	$ConvertNext_{i,t}$				$ConvertBack_{i,t}$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Shortfall	0.557		0.271		1.513		3.724***	
	(0.738)		(0.595)		(1.024)		(0.814)	
High Shortfall		-0.035		0.044		0.216		0.567**
		(0.235)		(0.210)		(0.262)		(0.223)
Board size	0.021	0.027	0.007	0.010	-0.141	-0.127	-0.112	-0.070
	(0.085)	(0.083)	(0.081)	(0.080)	(0.113)	(0.113)	(0.110)	(0.106)
Firm age	0.094	0.088	0.092	0.093	0.091	0.089	0.155	0.170
	(0.100)	(0.100)	(0.095)	(0.096)	(0.173)	(0.171)	(0.156)	(0.148)
Total assets	-0.133**	-0.132**	-0.120**	-0.121**	-0.194*	-0.188*	-0.228**	-0.224**
	(0.062)	(0.062)	(0.058)	(0.058)	(0.101)	(0.099)	(0.096)	(0.091)
ROA	-0.530*	-0.542*	-0.408	-0.412	-0.560	-0.591	-0.462	-0.522
	(0.296)	(0.298)	(0.287)	(0.288)	(0.404)	(0.398)	(0.394)	(0.379)
Leverage	0.427	0.453	0.351	0.361	1.003**	1.060**	0.928**	1.041**
	(0.313)	(0.308)	(0.292)	(0.290)	(0.459)	(0.455)	(0.426)	(0.418)
Largest owner	1.439***	1.366***	1.401***	1.385***	1.476**	1.348**	1.638***	1.331**
	(0.350)	(0.347)	(0.328)	(0.325)	(0.581)	(0.575)	(0.558)	(0.549)
Constant	-2.588***	-2.535***	-2.135***	-2.093***	-1.785*	-1.749*	-0.413	0.154
	(0.724)	(0.726)	(0.610)	(0.611)	(1.054)	(1.050)	(0.938)	(0.919)
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes		
Year fixed effects	Yes	Yes	No	No	Yes	Yes	No	No
Pseudo R^2	0.078	0.077	0.053	0.053	0.254	0.250	0.195	0.159
Firm-years	821	821	821	821	821	821	821	821

Table 8
Event-induced cumulative abnormal returns to equal-weighted portfolios of listed ASA

The abnormal return parameter for event k, AR_k, is estimated using daily portfolio stock returns converted to USD and industry-adjusted using the daily return on the FamaFrench FF49 industry portfolio. The exception is firms in *Offshore/Shipping*, which we industry-adjust using the portfolio return to 49 OSE-listed foreign firms in this sector. In Panel B, W^e is the daily return on the MSCI stock market world index in excess of the 3-month US T-bill. Since $d_{k,t}$ takes a value of one in the two-day event window (and zero otherwise), the reported two-day abnormal return is $CAR_k = 2AR_k$. Columns (2) and (3) uses a portfolio of *High shortfall* and *Low shortfall* firms, respectively, sorted at the year-end preceding the event date. The portfolio in column (4) is long in *High shortfall* firms and short in *Low shortfall* firms. *Cumulative* is estimated beginning prior to the first event and ending on the day of the last (fifth) event, with the dummy variable d redefined to take a value of one in each of the five two-day event windows, so that *Cumulative* $= 10AR$. p-values in squared brackets are based on the t-value for CAR_k, which is $t_k = AR_k/\sigma_{AR_k}$ (see the text for further details).

	All firms (1)	High shortfall (2)	Low shortfall (3)	High–Low (4)
A: Cumulative return ($2AR_k$) based on $r^e_t = \alpha + AR_k d_{k,t} + \varepsilon_t$				
February 22, 2002	-0.001	-0.007	0.004	-0.012
	[0.974]	[0.676]	[0.809]	[0.377]
March 8, 2002	0.020	0.019	0.021	-0.002
	[0.239]	[0.286]	[0.232]	[0.894]
June 16, 2003	-0.003	0.002	-0.010	0.012
	[0.882]	[0.942]	[0.680]	[0.556]
December 1, 2005	-0.007	-0.009	-0.005	-0.004
	[0.624]	[0.542]	[0.745]	[0.567]
December 9, 2005	0.010	0.008	0.012	-0.003
	[0.468]	[0.569]	[0.399]	[0.651]
Cumulative ($10AR$)	0.021	0.017	0.023	-0.007
	[0.595]	[0.687]	[0.597]	[0.845]
B: Cumulative return ($2AR_k$) based on $r^e_t = \alpha + AR_k d_{k,t} + \beta W^e_t + \varepsilon_t$				
February 22, 2002	0.000	-0.005	0.005	-0.011
	[0.983]	[0.738]	[0.733]	[0.402]
March 8, 2002	0.025	0.025	0.026	-0.000
	[0.114]	[0.105]	[0.105]	[0.969]
June 16, 2003	0.002	0.006	-0.003	0.010
	[0.894]	[0.761]	[0.873]	[0.625]
December 1, 2005	-0.005	-0.007	-0.003	-0.005
	[0.712]	[0.612]	[0.850]	[0.530]
December 9, 2005	0.011	0.009	0.013	-0.004
	[0.420]	[0.527]	[0.348]	[0.627]
Cumulative ($10AR$)	0.036	0.029	0.040	-0.011
	[0.319]	[0.461]	[0.304]	[0.732]

Table 9
Cross-sectional regressions for the event-induced cumulative abnormal returns

Coefficient estimates from cross-sectional OLS regressions for firm i and event k (listed in Table 2):

$$2AR_{i,k} = \alpha_k + \gamma_{1,k} Shortfall_{i,k} + \gamma_{2,k} \mathbf{X_{i,k}} + u_{i,k}, \quad i = 1, ..., N.$$

The dependent variable $2AR_{i,k}$ is firm i's two-day announcement-induced abnormal returns for event k, estimated as in Panel A of Table 8. The explanatory variables are *Shortfall* and the control variables in vector **X**. All variables are defined in Table 4 and measured at year-end prior to event k. Robust standard errors (White estimator) are reported in parenthesis. Significance levels: *** 1%, ** 5%, * 10%.

	Quota-related news event date k, ($k = 1, ..., 5$)				
	22-Feb-2002	8-Mar-2002	16-Jun-2003	1-Dec-2005	9-Dec-2005
Shortfall	-0.025	-0.003	0.003	-0.002	-0.004
	(0.061)	(0.058)	(0.060)	(0.019)	(0.022)
Total assets	0.008	0.002	-0.005	0.002	0.001
	(0.008)	(0.006)	(0.005)	(0.002)	(0.003)
Largest owner	-0.003	0.014	-0.010	-0.025*	0.015
	(0.026)	(0.037)	(0.038)	(0.014)	(0.014)
Government control	-0.018	-0.001	0.039	-0.002	-0.022*
	(0.032)	(0.029)	(0.025)	(0.010)	(0.011)
Codetermination	0.009	0.014	0.031*	-0.003	0.005
	(0.015)	(0.014)	(0.018)	(0.005)	(0.006)
Risk	-0.433	0.309	0.352	-0.449	0.153
	(0.371)	(0.304)	(0.566)	(0.296)	(0.739)
Constant	-0.058	-0.026	0.017	-0.011	-0.014
	(0.110)	(0.082)	(0.094)	(0.032)	(0.054)
R^2	0.10	0.02	0.07	0.09	0.02
Number of firms (N)	129	131	123	126	127

Table 10
Replicating Ahern and Dittmar (2012) with correction for stock return cross-dependence

Panel A lists the average cumulative abnormal returns (CAR^{AD}) reported by Ahern and Dittmar (2012) (AD) for N=94 OSE-listed firms over the five-day event-window (-2,2) around February 22, 2002, where

$$CAR_i^{AD}(-2,2) = \sum_{\tau=-2}^{2} (r_i - r_{imatch})_\tau \, .$$

Here, r_i is the return to OSE-listed ASA i, and r_{imatch} is the average return to US-listed companies in firm i's GICS industry. Return data are from Compustat Global (for Norwegian firms) and CRSP. AD report p-values only (in square brackets), so the corresponding standard error of the average $CAR^{AD}(-2,2)$ (in parentheses) is computed by us. Panel B shows our replication using the AD data and methodology. Panel C shows the portfolio estimate of the five-day abnormal return using the time series regression in Panel A of Table 8 but with the equal-weighted industry-adjusted portfolio return, $r_{p,t}^{-I} = \frac{1}{N}\sum_{i=1}^{N}(r_i - r_{imatch})_t$, as dependent variable (p-values in square brackets use the standard errors in the line above). Data from *Brønnøysund Register Centre* show that 69 firms have zero female directors in 2001 ($Zero_{2001}$), up from 68 in AD. Significance levels: *** 1%, ** 5% and * 10%.

Percent five-day abnormal return (CAR) centered on February 22, 2002	All firms in AD (1)	AD firms with $Zero_{2001}$ (2)	AD firms with Pos_{2001} (3)	Difference $Zero - Pos$ (2) - (3)
A: Original AD CAR estimates (no adjustment for cross-dependence of returns)				
Average $CAR^{AD}(-2,2)$	-2.573***	-3.547***	-0.024	-3.523***
St.err. of CAR^{AD} ($\sigma_{CAR} = \frac{\sigma}{\sqrt{N}}$)	(0.757)	(1.030)	(0.824)	(1.297)
p-value	[0.001]	[0.001]	[0.977]	[0.008]
Number of firms (N)	94	68	26	94
B: Replication of AD CAR estimates (no adjustment for cross-dependence of returns)				
Average $CAR^{AD}(-2,2)$	-2.733***	-3.738***	0.042	-3.780***
St.err. of CAR^{AD} ($\sigma_{CAR} = \frac{\sigma}{\sqrt{N}}$)	(0.780)	(0.973)	(1.011)	(1.403)
p-value	[0.001]	[0.000]	[0.967]	[0.009]
Number of firms (N)	94	69	25	94
C: Time-series estimation of CAR using $r_{p,t}^{-I} = \frac{1}{N}\sum_{i=1}^{N}(r_i - r_{imatch})_t = \alpha + ARd_t + \varepsilon_t$				
Average daily abnormal return over 5-day event window (-2,2)				
AR	-0.423	-0.661	0.210	-0.873
St.err. of AR (σ_{AR})	(0.650)	(0.685)	(0.711)	(0.563)
Five day CAR				
$CAR(-2,2) = 5AR$	-2.116	-3.305	1.051	-4.365
St.err. CAR ($\sigma_{CAR} = 5\sigma_{AR}$)	(3.250)	(3.435)	(3.555)	(2.815)
Statistical significance of CAR with adjustment for cross-dependence of returns				
p-value based on $t = \frac{CAR}{\sigma_{CAR}} = \frac{AR}{\sigma_{AR}}$	[0.516]	[0.336]	[0.768]	[0.122]
Number of firms (N)	94	69	25	94

Table 11
Long-run (36-month) abnormal stock performance of firms with all-male boards in 2001

Monthly abnormal stock returns for portfolios of listed ASA with zero ($Zero_{2001}$) and at least one (Pos_{2001}) female director in 2001. In Columns (3) and (6), the portfolio is long in $Zero_{2001}$ and short in Pos_{2001}. The estimation period is February 2002 (when the quota legislative process began) to April 2008 (when all firms complied). The monthly average number of firms in the two portfolios is 98 ($Zero_{2001}$) and 32 (Pos_{2001}). In columns (1)-(3), the abnormal stock return parameter α is estimated using the following three factor return-generating process:

$$r_t^e = \alpha + \beta_1 W_t^e + \beta_2 HML_t + \beta_3 SMB_t + \varepsilon_t,$$

where r_t^e is the monthly stock return to domestic OSE-listed ASA, converted to USD using the monthly exchange rate minus the monthly return on the firm's Fama-French FF49 industry portfolio. The exception is firms in the *Offshore/Shipping* sector, for which we subtract the monthly return to a value-weighted portfolio of 49 OSE-listed foreign firms in this sector. W^e is the monthly return on MSCI world stock market index in excess of the daily 3-month US Treasury bill. SMB (size) and HML (value) and, in columns (4)-(6), MOM (momentum) are global risk factors from Ken French's web site (*http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/*). Standard errors in parenthesis. Significance levels: *** 1%, ** 5%, * 10%.

Portfolio:	$Zero_{2001}$ (1)	Pos_{2001} (2)	*Zero-Pos* (3)	$Zero_{2001}$ (4)	Pos_{2001} (5)	*Zero-Pos* (6)
α	0.003	0.006	-0.003	0.003	0.006	-0.004
	(0.006)	(0.005)	(0.004)	(0.006)	(0.005)	(0.004)
W^e	0.155	0.316**	-0.161	0.169	0.311**	-0.142
	(0.142)	(0.131)	(0.100)	(0.152)	(0.139)	(0.107)
HML	0.742*	0.790**	-0.048	0.725*	0.795**	-0.070
	(0.394)	(0.362)	(0.279)	(0.401)	(0.369)	(0.283)
SMB	0.459	0.269	0.190	0.435	0.276	0.159
	(0.286)	(0.263)	(0.202)	(0.300)	(0.276)	(0.212)
MOM				0.049	-0.015	0.064
				(0.174)	(0.160)	(0.123)
R^2	0.112	0.133	0.050	0.113	0.134	0.053
Observations (months)	75	75	75	75	75	75

Table 12
Effect on Tobin's Q of appointing female directors to the board

Panel A reports estimates of the coefficient β from the second-stage instrumental variable (IV) regression:

$$Q_{i,t} = \alpha + \beta \widehat{Shortfall}_{i,t} + \theta_i + \tau_t + \epsilon_{i,t},$$

where θ and τ are, respectively, industry and year fixed effects. $\widehat{Shortfall}$ (plotted in Figure 3) is the fitted value from the first-stage IV regression (coefficients reported in Panel B):

$$Shortfall_{i,t} = \alpha + \sum_{\tau=T_1}^{T_2-1} \beta_\tau D_\tau Shortfall_{i,T_0} + \theta_i + \tau_t + u_{i,t}.$$

D_τ is a year dummy, and T_1 and T_2 are the beginning and ending years, respectively, of the estimation period. T_0 is the year in which the firm's exogenous quota-induced shortfall of female directors $Shortfall_{T_0}$ is measured. The sample comprises 239 OSE-listed ASA in 2002-2008 (columns (1) and (2)) and 227 OSE-listed ASA in 2003-2009 (columns (3)-(5)). Standard errors clustered by firm are reported in parenthesis. Significance levels: *** 1%, ** 5%, and * 10%.

Estimation period ($T_1 - T_2$):	2002-2008		2003-2009		
Year of exogenous shortfall (T_0):	2001	2000	2002	2001	2000
	(1)	(2)	(3)	(4)	(5)
A: 2^{nd} stage IV regression for Q					
$\widehat{Shortfall}$	0.750	0.535	1.910**	0.689	0.246
	(0.737)	(0.757)	(0.833)	(1.236)	(1.199)
F-statistic	18.526	19.754	15.851	16.691	17.224
Firm-years	815	726	820	790	683
B: 1^{st} stage IV regression for $Shortfall$					
$Shortfall_{T_0} \times D_{2002}$	0.714***	0.743***	.	.	.
	(0.096)	(0.102)	.	.	.
$Shortfall_{T_0} \times D_{2003}$	0.611***	0.655***	0.964***	0.634***	0.631***
	(0.129)	(0.121)	(0.091)	(0.131)	(0.130)
$Shortfall_{T_0} \times D_{2004}$	0.536***	0.637***	0.843***	0.553***	0.610***
	(0.114)	(0.096)	(0.108)	(0.120)	(0.115)
$Shortfall_{T_0} \times D_{2005}$	0.348***	0.439***	0.544***	0.360***	0.402***
	(0.105)	(0.103)	(0.108)	(0.116)	(0.122)
$Shortfall_{T_0} \times D_{2006}$	0.177**	0.268***	0.433***	0.184*	0.231**
	(0.081)	(0.088)	(0.096)	(0.095)	(0.100)
$Shortfall_{T_0} \times D_{2007}$	0.082*	0.046	0.167**	0.087	0.010
	(0.045)	(0.053)	(0.072)	(0.055)	(0.058)
$Shortfall_{T_0} \times D_{2008}$			0.058	0.006	-0.030
			(0.043)	(0.044)	(0.045)
F-statistic	84.789	76.731	85.885	45.545	40.387
Firm-years (N)	832	740	829	799	689

Table 13
Regressions for post-quota changes in operating profitability

The table shows coefficient estimates from panel OLS regressions of firm i's operating profitability $ROA_{i,t}$:

$$ROA_{i,t} = \alpha + \gamma_1 D^{Treat} D^{Comply}_{i,t} + \gamma_2 D^{Treat}_{i,t} + \gamma_3 D^{Comply}_{i,t} + \gamma_4 \mathbf{X}_{i,t} + \epsilon_{i,t}.$$

The treatment group is ASA, except in columns (2) and (6) in which it is ASA with pre-quota all-male boards ($Zero_{2001}$). D^{Comply} indicates year 2008 and onwards (when the quota is binding), except in columns (5)-(6), where it indicates years in which an ASA actually complies. The vector $\mathbf{X}_{i,t}$ contains the control variables listed below. All variables are defined in Table 4. The sample comprises 685 non-financial ASA and 2,627 non-financial Large AS, 2002-2013, and excludes AS registered as ASA at some point during the sample period. Standard errors clustered by firm are reported in parenthesis. Significance levels: *** 1%, ** 5%, and * 10%.

Compliance (D^{Comply}):	2008 and onwards				Actual	
Treatment (D^{Treat}):	ASA	$Zero_{2001}$	ASA		ASA	$Zero_{2001}$
Control (1-D^{Treat}):	Large AS	Pos_{2001}	Large AS		Large AS	Pos_{2001}
Estimation period	2002-2013	2002-2013	2003-2009	2003-2010	2003-2009	2003-2009
	(1)	(2)	(3)	(4)	(5)	(6)
$D^{Treat} D^{Comply}$	-0.009	-0.007	-0.034**	-0.021	-0.016	-0.002
	(0.013)	(0.031)	(0.015)	(0.013)	(0.014)	(0.039)
D^{Treat}	-0.146***	0.031	-0.143***	-0.141***	-0.144***	0.044
	(0.015)	(0.029)	(0.016)	(0.016)	(0.017)	(0.034)
D^{Comply}	-0.034***	-0.043*	-0.038***	-0.036***		-0.010
	(0.003)	(0.025)	(0.004)	(0.004)		(0.042)
Board size	-0.006***	-0.016**	-0.006***	-0.005***	-0.006***	-0.010
	(0.002)	(0.007)	(0.002)	(0.002)	(0.002)	(0.008)
Firm age	0.011***	0.027**	0.010***	0.011***	0.010***	0.015
	(0.002)	(0.011)	(0.003)	(0.003)	(0.003)	(0.011)
Total assets	0.028***	0.055***	0.027***	0.026***	0.028***	0.051***
	(0.003)	(0.007)	(0.003)	(0.003)	(0.003)	(0.007)
Leverage	-0.114***	-0.218***	-0.110***	-0.101***	-0.111***	-0.207***
	(0.019)	(0.045)	(0.024)	(0.022)	(0.024)	(0.056)
Largest owner	0.009	0.036	0.013	0.011	0.014	0.037
	(0.009)	(0.040)	(0.011)	(0.010)	(0.011)	(0.045)
Listed	0.009	-0.050**	0.025	0.018	0.024	0.042
	(0.018)	(0.025)	(0.020)	(0.019)	(0.020)	(0.029)
Constant	-0.150***	-0.566***	-0.148***	-0.142***	-0.144***	-0.444***
	(0.039)	(0.107)	(0.045)	(0.043)	(0.046)	(0.098)
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	No	No	No	No	Yes	Yes
R^2	0.188	0.232	0.191	0.183	0.192	0.222
Firm-years	12318	1919	7797	8847	7797	1307