Does gender-balancing the board reduce firm value?*

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Abstract

A board gender quota reduces firm value if it forces the appointment of under-qualified female directors. We test this hypothesis using Norway's 2005 board gender-quota law, which increased the average fraction of female directors from 5% in 2001 to 40% by 2008. Statistically robust analyses of quota-induced shareholder announcement returns, and of long-run stock and accounting performance, fail to reject the hypothesis of a zero valuation effect of this economy-wide shock to board composition and director independence. Evidence on female director turnover and changes in director networks also fails to suggest that qualified female directors were in short supply.

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1 Introduction

In December of 2005, Norway legally mandated its public limited liability companies (henceforth ASA) to have gender-balanced boards.[1] The quota requires 40% of shareholder-elected directors to be from each gender. ASA firms, about half of which are listed on the Oslo Stock Exchange (OSE), were given two years to comply or face liquidation—the ultimate penalty for violation of corporate law. This pioneering social experiment has generated substantial debate in corporate governance and labor market research, as well as among public policy makers. Board gender quotas have since been adopted by several European countries, beginning with Spain in 2007 and followed by Belgium, France, Germany, Iceland, Italy, and the Netherlands.[2]

The main purpose of this study is to test whether gender-balancing the board affects firm value. As such, our study adds to a growing literature using exogenous variation created by important legislative events to identify causal effects of changes in corporate governance on firm performance. The exogenous variation helps identify the direction of causality when underlying firm characteristics simultaneously determine optimal board composition, managerial incentives and firm performance (Bhagat and Jefferis, 2005; Adams, Hermalin, and Weisbach, 2010). For example, in the US, Bertrand and Mullainathan (2003) study the effect of state antitakeover law adoptions on managerial incentives and performance. Greenstone, Oyer, and Vissing-Jorgensen (2006) examine effects of early disclosure amendments to the 1934 Securities and Exchange Act, while Chhaochharia and Grinstein (2007) estimate effects of the 2002 Sarbanes-Oxley Act on firm value.

Since the incoming female directors are rarely part of any existing male director network, an immediate effect of the Norwegian gender quota is to increase director independence. While extant research indicates that increased director independence likely affects board decisions, the direction of the resulting valuation impact is unclear. For example, Adams, Hermalin, and Weisbach (2010) report that CEO turnover—a key board decision—is more sensitive to firm performance in companies with a high fraction of independent directors, which is likely positive for shareholders. On the other hand, there is evidence that gender-

[1] Private limited liability companies (AS) are not subject to the quota law. The ASA and AS legal forms reflect European Union standards and correspond, for example, to the PLC and Ltd forms in the UK, to S.A. and S.A.R.L. in France, and to AG and GmbH in Germany.

[2] The European Union (EU) is currently debating a gender-quota directive. During this change process, the average fraction of female directors in large publicly traded firms throughout EU has risen to 21% in 2015, driven primarily by quota countries. In the US, 18% of the directors of Fortune 1000 companies were female in 2015, up from 15% in 2011 (source: https://www.2020wob.com/companies/2020-gender-diversity-index). Resolutions urging board gender diversity for public companies have recently been passed in the California and Massachusetts state legislatures.

diverse boards tend to allocate more time to monitoring-type of board activities. Some studies argue this may have negative implications for shareholders (Adams and Ferreira, 2009; Bøhren and Staubo, 2015), while others argue that the greater monitoring is positive (Schwartz-Ziv, 2015).[3]

The key question for this paper is whether the quota-induced increase in director independence comes with sufficient female director qualifications and experience to avoid a reduction in firm value. We focus on the potential for value reduction because, as discussed below, extant empirical studies tend to report negative valuation effects. It is also possible for a gender quota to have positive valuation effects, e.g if it breaks an inefficient monopolization of the director election process by an "old boys" network. Or, the quota may have a value neutral effect if gender composition is largely irrelevant for firms' decision making. We test these mutually exclusive alternatives using short- and long-run performance metrics applied to the population of domestic and foreign companies listed on the OSE, as well as a large sample of non-listed firms.

To set the stage for our analysis, Figure 1 shows actual changes in the number of shareholder-elected directors in OSE-listed ASA over the sample period. Between 2005 and 2007—the formal quota-compliance period—the average fraction of female directors increased from 24% to 40%, which on average means replacing one male director with one female. When we also attribute some of the increase in female directors that begun in 2002 to shareholder anticipation of a future quota regulation, which we show later is a reasonable interpretation of the data, compliance means replacing closer to two of five male directors with females. Thus, the central empirical question is whether this level of forced board change significantly reduced the market value of firms traded on the OSE. This experimental setting is powerful both because the Norwegian quota was the first of its kind (and therefore relatively unanticipated by investors) and because it came with a particularly strong penalty for non-compliance (forced liquidation).[4]

Before turning to our empirical analysis, notice also in Figure 1 the near-constant average board size of five shareholder-elected directors over the sample period. Notwithstanding that small boards are probably more effective than large boards, shareholders who were sufficiently concerned with having to appoint potentially under-qualified female directors could have chosen to relax the quota constraint by

[3]Perhaps as a case in point: in 2002, the board of Statoil, Norway's largest publicly traded firm (a government-controlled oil company), was confronted with a bribery scandal involving a subcontractor in Iran. Two highly educated, newly appointed female directors refused to go along with a strategy of protecting the firm's insiders, ultimately forcing out the chairman and the CEO.

[4]Sanctions associated with quota laws subsequently adopted elsewhere in Europe vary considerably and are sometimes relatively weak. For example, Spain and the Netherlands have no sanctions, while non-compliance in German means leaving empty the board seats earmarked for women.

expanding board size. For example, under the quota law, a firm with a five-member all-male board could expand board size to eight in order to give room for three additional females without sacrificing valued male directors or board control. The fact that this did not happen (on average) is at least suggestive of a limited expected cost of the gender quota.

For direct valuation effects of the gender quota, we estimate announcement-induced abnormal stock returns to both domestic and foreign OSE-listed companies, where the latter are not subject to the gender quota. Second, we complement the event-study analysis with panel estimation of Tobin's Q as well as long-run abnormal stock return and accounting performance, where the latter covers non-listed firms as well. For indirect valuation effects, we examine the extent of corporate conversions from ASA to the AS legal form in order to circumvent the quota. Finally, we present new large-sample evidence on the evolution of board seat dispersion and director turnover and experience, which helps assess the depth of the pool of qualified female directors.

Our short-run, event-study analysis identifies eleven key quota-related news events over the lengthy time period that led to the quota being signed into corporate law (from August 1999 through December 2005). Using a robust estimation methodology that includes both domestic and foreign firms listed on the OSE, we cannot reject the hypothesis of a value-neutral effect of these news announcements, either individually or combined. Furthermore, in the long-run performance analysis, we find no traces of significant risk-adjusted portfolio stock returns in calendar time, nor an effect on Tobin's Q of the shortfall female directors caused by the quota. Moreover, our estimates of long-run operating profitability further supports the market value analysis as we find no statistically significant drop in profitability attributable to the new female directors over the post-quota sample period (2009-2013).

Our work on short- and long-run valuation effects is related to several studies in the literature that reach different conclusions. Nygaard (2011) presents evidence of significantly *positive* average abnormal stock return to OSE-listed companies when the quota was signed into law in December of 2005. However, we show that foreign firm listed on the OSE also experience positive abnormal returns of a statistically similar magnitude relative to that event, suggesting that the abnormal returns arise for reasons other than the quota constraint on domestic firms. Ahern and Dittmar (2012) report significantly *negative* abnormal stock returns centered on an event in February of 2002 that also increased the probability of a quota law. However, we show that their event analysis straddles two opposing quota-related news events. Separating the two events leads to statistically insignificant abnormal stock returns also when we otherwise replicate their sample and estimation procedure.

The result of our long-run performance analysis also contrasts with that of Ahern and Dittmar (2012). They conclude that the gender quota imposed a strongly binding constraint, suggesting value losses of upwards of 20% for firms with no prior female directors.[5] However, we show that this conclusion hinges both statistically and economically on their instrumentation. Finally, our finding of a statistically insignificant accounting performance differs from Matsa and Miller (2013), who provide evidence of an increase in labor costs and decrease in operating profitability. Our longer post-quota sample period (through 2013) produces results indicating that the negative accounting performance in Matsa and Miller (2013) is driven by the financial crisis years rather than the 2005 gender quota.

As pointed out by Nygaard (2011) and Bøhren and Staubo (2014), the rate of ASA to AS conversions may also provide information on the perceived cost of the quota restriction. Over the period 2001 through 2007, no OSE-listed ASA converted to AS except in the event that the firm was liquidated or acquired—events that are unrelated to the quota restriction. Thus, whatever the perceived cost of the gender quota, it was apparently lower than the OSE listing benefit *per se*. As for *unlisted* ASA, the population began to decline after peaking in 2001. However, contrary to the conclusion of Bøhren and Staubo (2014), our panel estimation shows that this decline is statistically independent of the quota-induced female director shortfall, and thus cannot be attributed to the quota restriction.

We end the paper by providing large-sample descriptive information on the evolution of board seat dispersion and director turnover and experience. Such evidence may uncover firm responses to the quota restriction not necessarily detected by the direct valuation analysis or ASA to AS conversions. In particular, if qualified females are in short supply, we expect board seats to become concentrated among a few but "busy" female directors. Moreover, director turnover among female directors would be relatively high as boards attempt to replace under-qualified female directors *ex post*.

However, these two final predictions related to a negative valuation impact of the quota are also not supported by our data. First, director seat dispersion is wide for both men and women, and has remained largely unaffected by the quota. Of the female directors in listed ASA, only 6% hold three or more board seats (our definition of busy director) over our six-year sample period following mandatory quota compliance (2008-2013), while the corresponding percentage for male directors is 2%. This is

[5] To place their conclusion in perspective, value losses upwards of 20% are typically reserved for dramatic corporate events such as product recalls (Jarrell and Peltzman, 1985), corporate bankruptcy and distressed restructurings (Hotchkiss, John, Mooradian, and Thorburn, 2008).

substantially less than the 25% busy directors among S&P 1500 companies over the period 1999-2008 (Cashman, Gillan, and Jun, 2012). Second, consistent with a sufficiently deep pool of qualified director candidates, there is also no evidence that female director turnover on average increases post-quota. Finally, although incoming female directors have far less CEO experience than their male counterparts (Ahern and Dittmar, 2012), boards appear to have maintained a relatively constant overall director CEO experience. This is consistent with a strategy of replacing the least experienced male directors, and with an increasing pool of qualified female directors over time (Bertrand, Black, Jensen, and Lleras-Muney, 2014).

In sum, the paper demonstrates the following six empirical regularities: (1) Insignificant abnormal stock and accounting performance (both short- and long-run), (2) constant board size over the sample period, (3) ASA to AS conversions that are statistically independent of the quota-induced female director shortfall, (4) wide male and female board seat dispersion both before and after the quota, (5) normal post-quota female director turnover, and (6) relatively constant overall director CEO experience for ASA boards. Collectively, this evidence suggests that the Norwegian gender-quota was perceived to have (and indeed had) a value-neutral effect on firm value, and that shareholder-borne costs associated with quota were modest.

The rest of the paper is organized as follows. Section 2 estimates the market reaction to quota-related events, while Section 3 presents a long-run stock and accounting performance analysis. Section 4 examines ASA to AS conversions, and section 5 presents our descriptive evidence on seat dispersion and director turnover. Section 6 concludes the paper.

2 Market reaction to quota news events

The Norwegian gender quota was signed into law in December of 2005 and firms were given until the end of 2007 to comply. While Norway's codetermination laws grant employees certain rights to elect directors among themselves, the quota law itself regulates ASA-directors elected by shareholders only. Table 1 shows how the mandated fraction of females varies with the shareholder-elected board size. In boards with three directors, one director (33%) must be of each gender. In boards with four or more shareholder-elected directors, the mandated percentage female (and male) directors ranges from 38% to 50%.

A gender quota may reduce firm value if qualified female directors are in short supply, forcing the appointment of under-qualified women to boards. To examine this potential valuation effect, we begin with an estimation of abnormal stock returns in response to key quota-related news announcements for the population of domestic and foreign OSE-listed firms, 1998-2005. As the quota applies to domestic ASA only, we use the foreign OSE-listed firms as a control group in the event study.

2.1 Key quota-related news announcements

We start by a complete mapping of all important quota-related news events ending with the signing of the quota amendment into corporate law in December of 2005. As summarized in Table 2, there are 11 key news announcements, beginning in 1999 with an initial discussion of an amendment to Norway's 1978 gender equality law (event 1, Panel A). Section 21 of this law (in a 1981 amendment) requires at least 40% of members on government-appointed committees to be from each gender. This is the intellectual origin of the 40% rule embedded in the subsequent corporate quota (Table 1).

The debate switched in 2001 to one of amending corporate law (Panel B), championed by the then minority center-left government. However, parliamentary elections in September 2001 brought in a center-right coalition government opposed to a gender quota for corporate boards. Thus, at the end of 2001, the probability of a gender quota law was slim. Then, in a surprise move, the Minister of Trade and Industry signaled, through an interview in a tabloid newspaper on February 22, 2002, that he supported a quota (event 5 in Panel C). The next day, however, he retracted his quota support in the daily business newspaper (event 6). This was followed by a unanimous rejection of a quota by the parliamentary members of his conservative party (the largest political party in the center-right coalition). However, following negotiations within the center-right coalition (led by Laila Dåvøy, the Minister of Children and Family Affairs) the Cabinet made a surprise announcement on March 8, proposing quota legislation (event 7).

The March 8, 2002 proposal announcement included a promise that the quota would be cancelled if firms voluntarily gender-balanced their boards before the end of 2005. Thus, the government's intention was to induce voluntary corporate action. At the same time, the government stated that it would rapidly gender-balance government-owned firms. This political objective appears to have affected the five government-controlled OSE-listed firms in our study. By the end of 2002, these firms had on average

39% female directors, up from 31% the year before.[6]

The Norwegian Parliament passed the proposed amendment in November of 2003 (event 9). The amendment contained an important provision: the quota would not be legislated if firms complied voluntarily within two years. Voluntary compliance did not occur, however, and the Cabinet signed the board gender quota into law on December 9, 2005 (event 11, Panel D). The December 9 announcement also resolved uncertainty about the penalty for non-compliance. As recent as December 1 (event 10), the Prime Minister appears to confirm market expectations that the penalty would be restricted to a monetary fine. However, the penalty implied by the amendment enacted on December 9 is forced liquidation—the ultimate penalty for violations of corporate law. Firms were given two years to comply and by year-end 2007, the average OSE-listed ASA had 39% female directors.[7]

The news events in Table 2 confirm that 2001 is the last year in which the probability of a gender quota is negligible for valuation purposes. New announcements, starting February 22, 2002, began to increase the probability of a quota. In the event-study analysis below, we estimate the market reaction to all eleven events and make inferences based on the statistical significance of the events both individually and jointly. This ensures that our overall inferences are robust to variations in the level of news content across the eleven announcements.

2.2 Data and estimation methodology

We obtain stock price data from *Børsprosjektet* at the Norwegian School of Economics, whose direct data source is the OSE data service *Oslo Børsinformasjon*. This data source contains stock prices, dividends and number of shares for all OSE-listed companies (beginning in 1967). 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, which occurs in twenty percent of the trading days.

Since the quota events affect all firms simultaneously, we form an equal-weighted portfolio of the OSE-listed firms in calendar time. Panel A of Figure 2 shows the number of OSE-listed companies from 1998 through 2013. There are 339 unique OSE-listed firms, of which 290 are domestic and 49 are foreign

[6] Telenor ASA (telecommunications), Norsk Hydro ASA (oil and gas), Statoil ASA (oil and gas), DNB ASA (bank), and Kongsberg ASA (weapons).

[7] Of the listed ASA, 86% were in compliance by year-end 2007, and 100% of all (listed and unlisted) ASA had complied by April 2008.

companies. These 339 unique firms enter and exit the OSE at different points over the sample period. Panel B shows the number of firms in the portfolio through December 9, 2005 (the date the quota was signed into law), which we use for the abnormal return estimation described below. The number of firms in the domestic portfolio average 150 (ranging from 128 to 171), while there are on average 22 firms in the foreign portfolio (ranging from 14 to 29).

Daily abnormal portfolio stock returns are estimated separately for each of the eleven events in Table 2, using stock returns from day -251 through day 0 (the event day). For a firm to be included in this estimation, it must have return observations on *both* days in the two-day event window (-1,0) and at least one hundred return observations in the estimation period (-251, -2).

We use three model specifications to estimate abnormal stock returns. In the first model, we simply compute the mean-adjusted return (no explicit risk-factor adjustment), which minimizes uncertainty about parameter estimation. This mean-adjusted return is the parameter AR_k in the following event-time regression, where day 0 is the day of the k^{th} news event:

$$r^e_{k,t} = \alpha_k + AR_k d_{k,t} + \varepsilon_{k,t}, \quad t = -251, .., 0. \tag{1}$$

Here, $r^e_{k,t}$ is the daily, equal-weighted portfolio return in excess of the one-day Norwegian interbank offer rate (NIBOR, obtained from the Norwegian Central Bank), and $d_{k,t}$ is a dummy variable, taking a value of one in the two-day window (-1, 0) of event k and zero otherwise. The parameter AR_k measures the average daily abnormal return over the event window, so the two-day event-window abnormal return is $2AR_k$.[8]

The estimation of AR_k is performed for each of the eleven events in Table 2, always ending the estimation period on day 0. Whenever the estimation period for the k^{th} event includes one or more prior events in Table 2, we add a dummy variable to equation (1) for each of the prior two-day event windows, which removes the influence of these prior event returns on the estimate of AR_k.

Our second approach uses the following market model with a lagged world market index:

$$r^e_{k,t} = \alpha_k + AR_k d_{k,t} + \beta_{k1} W^e_t + \beta_{k2} W^e_{t-1} + \varepsilon_{k,t}, \quad t = -251, .., 0. \tag{2}$$

[8] Since event date (6) is a Saturday, day 0 is the following Monday (the first trading day). Moreover, since day -1 is the event day for event (5), we use the window (0,1) to estimate abnormal return for event 6.

Here, W_t^e is the contemporaneous daily excess return on the Morgan Stanley Capital International (MSCI) world stock market index, converted to NOK using the daily exchange rate, and in excess of NIBOR. MSCI is from Thomson-Reuters and the NOK/USD daily exchange rate is from the Norwegian Central Bank. The regression model includes the lagged market return, W_{t-1}^e, because there is substantial evidence that the OSE market is predictable by this lagged world market return (Eckbo and Smith, 1998; Næs, Skjeltorp, and Ødegaard, 2009).[9]

The third model, for robustness, estimate abnormal returns with the following expanded return generating process:

$$r_{k,t}^e = \alpha_k + AR_k d_{k,t} + \beta_{k1} W_t^e + \beta_{k2} W_{t-1}^e + \beta_{k3} HML_t + \beta_{k4} SMB_t + \beta_{k5} MOM_t + \varepsilon_{k,t}, \tag{3}$$

where HML and SMB are the value and size factors (Fama and French, 1992) and MOM is the momentum factor (Carhart, 1997). These factors, which are calculated using the population of OSE stocks, are made available by Bernt A. Odegaard on his web site, http://finance.bi.no/∼bernt/.

2.3 Average abnormal stock returns

Panel A of Table 3 and of Table 4 lists the portfolio abnormal return estimates for each of the eleven quota-related news events in Table 2. In the following we focus on the results in Table 3, and treat Table 4 as robustness. The first three columns in Table 3 report results using the mean-adjusted model in equation (1), while the next three columns use the two-factor model in equation (2). Panel B sums the abnormal returns over the events in each of the four legislative subperiods listed in Panels A through D of Table 2. In the following discussion, in order to conclude that the gender quota affects firm value, we require evidence of significant abnormal returns to domestic firms *and* to the long-short portfolio of domestic minus foreign firms. The latter is necessary in order to infer that the abnormal return is driven by the quota *per se*.

Beginning with the portfolio of domestic firms in column (1) and Panel A of Table 3, three of the eleven events produce statistically significant abnormal returns, of which one is significantly negative (event (8)). Recall that on this event date, June 13, 2003, the Cabinet submits the quota proposal

[9] As discussed by Thompson (1985), the conditional event-parameter estimation in regression equation (2) is equivalent to a traditional residual analysis (MacKinlay, 1997) whenever the event is uncorrelated with the market return, which is the case in our study of involuntary corporate events (Eckbo, Maksimovic, and Williams, 1990).

to Parliament. However, as shown in column (4), the negative abnormal return becomes insignificant when adjusting for the market risk factors in equation (2). Moreover, regardless of the return model, the abnormal stock return to the long-short portfolio is statistically insignificant. Thus, the abnormal returns associated with event (8) fail to support the hypothesis that the quota law affects firm value.

We next turn to the two events with positive abnormal returns in column (1). Recall that on March 8, 2002 (event 7), the Cabinet made a surprising turnaround and announced its decision to introduce a quota. The abnormal return of about 3% around this event is positive and statistically significant for domestic firms regardless of the return model, suggesting that the quota has a positive effect on firm value. However, this conclusion is premature as the news hitting the market on March 8, 2002, also create significantly positive abnormal returns to the portfolio of foreign firms, with statistically insignificant abnormal returns to the long-short portfolio. Thus, we conclude that the positive news on March 8, 2002, driving the positive abnormal stock returns, are unrelated to the quota.

Third, Table 3 shows evidence of positive abnormal stock returns of about 1% over event window (10). Recall from Table 2 that, on December 1, 2005, the Prime Minister confirms in an interview that the Cabinet will sign the gender quota into law as voluntary compliance did not happen in time. The abnormal return to foreign firms is statistically insignificant for this event, and the long-short portfolio abnormal return is a positive 1.2% and weakly significant at the 5% level. This evidence rejects the hypothesis of a negative valuation effect of the quota, and give some support to the hypothesis of a positive valuation effect.

There is a chance that one or more of the eleven events in Table 3 will emerge randomly as statistically significant even if the true abnormal returns are zero. Because of this, the 5% level of significance of the abnormal return to the long-short portfolio around event (10) may be overstated. To address this issue, Panel B of Table 3 reports the average cumulative abnormal returns across all events within each of the four legislative sub-periods in Table 2. That is, in the estimation of the return generating process (2) above, subscript k now refers to the k'th legislative period, and the event dummy $d_{k,t}$ takes a value of one in all of the two-day event windows covered by the subperiod. Thus, for each legislative period, the cumulative abnormal return reported in Panel B of Table 3 is $CAR_k \equiv K_k \times AR_k$, where K_k is twice the number of events in the k'th subperiod ($k = 1, .., 4$). The estimation periods starts 251 trading days prior to the first event in the subperiod and ends on day 0 of the last event.

In columns (1) and (4), abnormal returns summed across each of the first three legislative periods are statistically insignificant. Moreover, abnormal returns summed over the last two events (the last row of Panel B) is 1.5% weakly positive at the 5% level. However, as shown by the long-short portfolio in columns (3) and (6), this abnormal return is indistinguishable from that of the abnormal return to the portfolio foreign firms. In sum, the evidence in Table 3 fails to reject the null hypothesis of a value-neutral impact of the eleven quota news events.

The above analysis also has implication for previous event studies on the Norwegian quota. First, Nygaard (2011) finds significantly positive average abnormal returns to domestic OSE-listed firms around December 9, 2005 (our event 11) using a single-factor model with the contemporaneous world market index (MSCI). This is consistent with our abnormal return estimate for event (11) when we use the multifactor return model in Table 4, but not when we use the two simpler models in Table 3. However, Column (3) of Panel A of Table 4 also shows that the abnormal returns to domestic and foreign firms are statistically indistinguishable. Since foreign firms are not subject to the gender quota, we conclude from this that the information behind the positive market reaction on December 9, 2005, is most likely unrelated to the quota *per se*.

Moreover, Ahern and Dittmar (2012) focus their event study on Friday February 22, 2002 (our event 5). They use a five-day event window (-2,2), which stretches from Wednesday through Tuesday the following week. Recall from Table 2 that the Minister of Trade who announced his support for a quota on Friday, publicly retracted his support the next day (our event 6). As shown in Appendix Table 1, when we split up the two adjacent events 5 and 6, but otherwise replicate Ahern and Dittmar (2012)'s sample and abnormal return estimation, their abnormal returns become statistically insignificant and of a similar magnitude (-0.8%) to that reported in Panel A Table 3 above.[10]

2.4 Determinants of abnormal returns over legislative subperiods

While the above analysis fails to identify significant event-induced abnormal returns to portfolios of OSE-listed firms, Table 5 examines whether there is systematic cross-sectional variation in the firm-specific abnormal return estimates cumulated across legislative subperiods. Thus, in the cross-sectional regression, the dependent variable for firm i over the k'th legislative period is $CAR_{i,k} \equiv K_k \times AR_{i,k}$. As

[10]We thank Kenneth Ahern and Amy Dittmar for supplying the Compustat identifiers for their event-study sample of 94 OSE-listed firms. In Appendix Table 1, we use their data sources for all information except for the number of female directors and board size at year-end 2001, which is based the *Brønnøysund* register, obtained electronically through the Norwegian School of Economics (Berner, Mjøs, and Olving, 2013).

for the portfolio estimate of CAR_k in column (4) of Table 3 above, $CAR_{i,k}$ is estimated using Eq. (2) as the return generating process. The following cross-sectional regression is estimated using OLS:

$$CAR_{i,k} = \alpha_k + \beta_k \mathbf{X_{i,k}} + u_{i,k}. \tag{4}$$

The vector $\mathbf{X}$ contains six firm characteristics, collected from the national corporate registry at year-end prior to the year of the first event in the legislative subperiod: *Shortfall women* is the quota-induced shortfall of female directors, computed as the maximum of zero and the difference between the exact quota requirement (column (3) of Table 1) and the actual fraction of female directors. If the gender quota causes high-quality male directors to be replaced by lower-quality females, *Shortfall women* is predicted to have a negative effect.

Firm size is the natural logarithm of the market value of equity. Firm size is expected to have a negative impact on the announcement returns because small growth companies, which often rely on board advice (Field, Lowry, and Mkrtchyan, 2013), may be at a disadvantage in the competition for high-quality female directors. Furthermore, *Codetermination* is a dummy indicating that the female directors required by the quota and the employee directors together hold at least half of the board seats. We include this variable to account for the possibility that employee-elected directors form coalitions with the incoming females to try to lower the risk of the firm (Loderer, Benelli, and Lys, 1987). Hence, we also control for daily stock return volatility (*Risk*), measured over the associated event window. If voting coalitions involving employee-elected directors potentially reduce firm risk, the coefficients for *Codetermination* and *Risk* are predicted to be negative.

Finally, *Government control* is a dummy variable taking the value of one for the five government-controlled firms listed in footnote 6 above and *Largest owner* is the percent equity ownership of the largest shareholder. Since the incentive to monitor the board increases with shareholder concentration, *Largest owner* is predicted to have a positive effect on announcement returns. The regressions also include industry sector dummies.

The regression results are shown in Table 5. *Shortfall women* is uniformly insignificant, again suggesting that the overall economic effect of the gender quota is value neutral. The other individual coefficient estimates are also for the most part statistically insignificant. The exception is *Firm size*, which is significant at the 1% level in legislative period B. However, the significance of this variable is generally weak as it produces an insignificant coefficient in periods C and D, when the most important steps towards the quota were taken.[11]

The above firm-level event-study analysis unambiguously fails to reject the hypothesis of a value-neutral effect of news events up through December 9, 2005, including events that may have significantly increased the probability of a gender quota. We next turn to an estimation of potential long-run valuation effects beyond 2005.

3 Gender balancing and long-run performance

The event study analysis estimates the market reaction to news affecting the probability of a gender quota. This market reaction reflects the expected quality and impact of the new female directors forced by the quota. In this section, we examine long-run stock and accounting performance as actual board composition changes over the sample period. Actual board changes that deviate from market expectations may give rise to abnormal stock returns, which we estimate below using both monthly and annual data. The calendar-time stock portfolios end in April of 2008, when all ASA had fully complied with the law. The accounting performance analysis extends through 2013.

3.1 Performance of long-short stock portfolios sorted on female directors

Recall from Table 2 that the probability of a gender quota being legislated was significantly affected by events starting in February of 2002. Figure 3 shows the cumulative monthly (raw) returns to two equal-weighted portfolios of OSE-listed firms sorted on their female directors in 2001. One portfolio contains the firms with at least one female director by year-end 2001 (low shortfall) and the other portfolio contains the firms with zero female directors by year-end 2001 (high shortfall). Firms never switch portfolios and they enter if and when they list and exit if and when they delist.

The monthly average number of firms in the low-shortfall and high-shortfall portfolios is 32 and 98, respectively. The return cumulation starts in February of 2002—the beginning of significant public discussion of a quota law—and ends in April of 2008, when all firms are in full compliance. The figure

[11]The conclusions from Table 5 are robust to using abnormal returns to each of the 11 events individually as dependent variable and to the choice of return-generating process. Moreover, if we replace *Shortfall women* with a dummy variable indicating that there is no female director at the last year-end prior to each respective event, then this dummy also receives a statistically insignificant coefficient.

also plots the monthly cumulative return on the MSCI world stock market index, converted to NOK.[12]

In Figure 3, notice first the dramatic drop in value for both portfolios through early 2003. This drop parallels that of the MSCI world index following the collapse of tech sector stocks (beginning in year 2000), and which is an important driver of OSE market returns as well. The value drop in 2002 was greater for the high-shortfall portfolio than the low-shortfall portfolio. However, since the event study above provides no evidence that firms were affected negatively by news increasing the probability of a quota, and since there is no difference in announcement returns between high- and low-shortfall firms (Table 5), it is highly unlikely that the value drop is related to the fraction of female directors for either of the two portfolio.

If the forced addition of female directors up until April 2008 is truly costly for shareholders, then firms with zero female directors in 2001 have the most to loose from the quota law. Moreover, significant losses would then lead to under-performance relative to firms already (voluntarily) having female directors, with the under-performance gradually cumulating over time until April of 2008. However, Figure 3 does not indicate that the high-shortfall portfolio under-performs the low-shortfall portfolio after the market turned in 2003.

Table 6 shows the results of a formal performance analysis of the portfolios in Figure 3. The full pricing model, reported in columns (4)-(6), is:

$$r_t^e = \alpha + \beta_1 W_t^e + \beta_2 W_{t-1}^e + \beta_3 HML_t + \beta_4 SMB_t + \beta_5 MOM_t + \varepsilon_t, \tag{5}$$

where, as before, supercript e indicates return in excess of NIBOR, and HML, SMB and MOM are the value, size and momentum factors, respectively, calculated for the population of OSE stocks (see footnote 9 above for data sources). The first three columns of Table 6 restrict the risk model to the contemporaneous (W_t^e) and lagged (W_{t-1}^e) values of the MSCI world market index (in NOK). The calendar time estimation starts in February of 2002 and ends in April of 2008—a total of seventy-five months.

The evidence in Table 6 provides no support for the hypothesis that the quota law is costly for shareholders. Confirming the impression given by Figure 3, in columns (3) and (6), where the portfolio is long in firms with low shortfall and short in firms with high shortfall female directors in 2001, the abnormal performance measure α is statistically insignificant. Moreover, the estimate of α for the high-

[12]The NOK/USD exchange rate fell from 9.01 on December 31, 2001, to 5.13 on April 30, 2008, thus lowering the MSCI world market return expressed in NOK. Note also that OSE itself has a $\beta > 1$ against the MSCI.

shortfall portfolio is zero (columns (2) and (5)). While α is positive and significant at the 5% and 1% level in columns (1) and (4), respectively, this is either because firms with low shortfall truly outperform the market, or because the pricing model in equation (5) omits true risk factors. Note that the α of the long-short portfolio in columns (3) and (6), which is insignificant, accounts for omitted risk factors as long as the associated factor exposures are identical across the two portfolios.

3.2 Female director shortfall and Tobin's Q

In this section, we examine whether exogenous variation in *Shortfall women* affects Tobin's Q. We define Q as (book value of total assets − book value of equity + market value of equity)/book value of total assets. The data used to compute Q is from *Børsprosjektet*.[13] As in the cross-sectional analysis of event-induced abnormal returns in Section 2.4 above, we use the pre-quota $Shortfall\ women$ to identify exogenous variation in the firm-specific costs of complying with the quota requirement. Given the important quota-related news events in 2002 (Table 2), which were followed by an increase in female board representation for 29% of the OSE-listed ASA from year-end 2001 to year-end 2002, "pre-quota" means prior to 2002.

Before implementing the two-stage IV procedure in Panel B and C, Panel A of Table 7 shows the estimated slope coefficient on *Shortfall women* when using the following reduced-form, fixed effects panel regression:

$$Q_{i,t} = \alpha + \beta Shortfall\ women_{i,t} + \theta_i + \tau_t + \epsilon_{i,t}, \quad t = T_1, .., T_2, \tag{6}$$

where T_1 and T_2 are the beginning and ending years of the regression period, and θ_i and τ_t are firm and year fixed effects. Here, the estimation starts with *Shortfall women* in year $T_1 = 2002$ and ends in Column (1) with the formal $T_2 = 2007$ deadline for complying with the quota law (a panel of 847 firm-years). Column (1) in Panel A shows that Q is statistically unrelated to $Shortfall\ women$, with an estimated slope coefficient of 0.07 (clustered standard error of 0.34). In Column (2), we extend the estimation period to $T_2 = 2008$ because 14% of the OSE-listed firms did not comply until early that year. The estimated slope coefficient for the extended sample period, which comprises 1,009 firm-years, is 0.30 and again statistically insignificant.

In Panel B and C of Table 7, we use the pre-quota variation in *Shortfall women* to identify the

[13] The market value of equity is stock price times shares outstanding (shares issued - treasury shares), using the last closing price of the year. If a firm has more than one equity class (i.e. both A and B shares), the market value of equity is the combined market value of all equity classes. We drop Tobin's Q values that are less than or equal to zero, and then winsorize the resulting observations at 1% and 99% each year.

exogenous variation in $Shortfall\ women_{i,t}$ over the full compliance period, from $T_1 = 2002$ through $T_2 = 2007$ (column 1) and, alternatively, $T_2 = 2008$ (column 2). Building on Stevenson (2010) and Ahern and Dittmar (2012), the idea is to remove the time series variation in $Shortfall\ women_t$ that is self-selected by the firm in order to ultimately comply with the quota. Specifically, in the first stage, we regress $Shortfall\ women_t$ on $Shortfall\ women_{T_1-1}$ interacted with year dummies, as follows:

$$Shortfall\ women_{i,t} = \alpha + \sum_{\tau=T_1+1}^{T_2} \beta_\tau D_\tau Shortfall\ women_{i,T_1-1} + \theta_i + \tau_t + u_{i,t}, \quad t = T_1, .., T_2, \tag{7}$$

where D_τ is a year dummy. The fitted value of this regression forms the instrument $\widehat{Shortfall\ women}_t$ used in the second stage. Essentially, given $Shortfall\ women_{2001}$, the regression generates a predicted value $\widehat{Shortfall\ women}_t$ in each year from $T_1 = 2002$ until the deadline for mandatory compliance (T_2), at which point the shortfall equals zero for all sample firms.

In the second stage, we estimate the following OLS regression:

$$Q_{i,t} = \alpha + \beta \widehat{Shortfall\ women}_{i,t} + \theta_i + \tau_t + \epsilon_{i,t}, \quad t = T_1, .., T_2. \tag{8}$$

Panel B of Table 7 reports the estimate of the slope coefficient β. As shown, Q is statistically unrelated to $\widehat{Shortfall\ women}$ in the regression. Specifically, the slope-coefficient estimate is 0.80 when the estimation period ends in $T_2 = 2007$ (Column 1) and 0.75 when it ends in $T_2 = 2008$ (Column 2). Thus, consistent with the event study analysis above, we find no statistically significant relationship between the shortfall female directors and Q, whether or not instrumented.[14]

Ahern and Dittmar (2012), while using an econometric structure that is very similar to ours in Table 7, reach a substantially different conclusion. As shown in Table 8, this difference is largely driven by Ahern and Dittmar (2012)'s choice of year-end 2002 to identify the starting values of the percentage female directors used in the first-stage instrumentation. Table 8 repeats the regressions in Table 7 using $Shortfall\ women_{2002}$ in the first-stage estimation and with $T_1 = 2003$ and $T_2 = 2009$ (as in Ahern-Dittmar). Panel B now reports a second-stage slope coefficient on $\widehat{Shortfall\ women}$ that is statistically significant at the 5% level in column (3) (and also in Column 2, with $T_2 = 2008$). The coefficient estimate in Column (3) is 1.91 which is almost identical to the estimate -1.92 reported by Ahern and Dittmar

[14]This conclusion is unchanged if we also include firm-specific control variables, such as the natural logarithm of book value of assets, book leverage, and board size.

(2012).[15]

It is not surprising that using $Shortfall\ women_{2002}$ in the first-stage regression should matter so much for the test outcome. Recall from Section 2.1 that 29% of the OSE-listed ASA increased their female board representation from year-end 2001 to year-end 2002 following the several important quota events that year (Table 2). Recall also that the five large government-controlled sample firms had nearly fully complied with the (subsequent) quota at the end of 2002. In other words, the cross-sectional distribution of the percentage female directors at year-end 2002 reflects firm-specific efforts to comply with the quota ahead of time. As a result, $Shortfall\ women_{2002}$ violates the exclusion requirement for proper instrumentation. Panel C and D of Table 8 show that replacing $Shortfall\ women_{2002}$ with either $Shortfall\ women_{2001}$ or $Shortfall\ women_{2000}$ in the first stage again produces a small and statistically insignificant slope coefficient in the second stage.[16]

3.3 Gender balancing and operating profitability

In this section, we ask whether the quota affects operating profitability, measured as return on assets (ROA), where ROA=EBIT/Total assets. This serves to complement the above market-value based analysis. We estimate the following panel regression for firm i in year t:

$$ROA_{i,t} = \alpha + \gamma_1 ASA \times Post_{i,t} + \gamma_2 Post_{i,t} + \gamma_3 ASA_{i,t} + \gamma_4 \mathbf{X}_{i,t} + \epsilon_{i,t}. \tag{9}$$

The dummy variable $Post$ takes the value of one either in years after 2006 ($Post2006$) or 2008 ($Post2008$). The former starts the post-quota period in 2007, while the latter starts the post-quota period in 2009. As discussed above, since 2009 is the first full year in which all ASA complied with the quota, $Post2008$ more accurately reflects the post-quota decisions taken by gender-balanced boards. However, as discussed below, we include $Post2006$ in order to better compare our findings to those of Matsa and Miller (2013).

[15] The sign of the coefficient is switched because we use shortfall women based directly on the quota restrictions in Table 1, while Ahern and Dittmar (2012) use the near-inverse variable: percentage of women directors. There are also some other minor differences between the two studies, none of which affect the above conclusion: Ahern and Dittmar (2012) use industry-adjusted Q while we use Q itself (Gormley and Matsa, 2014); they have a somewhat smaller sample size (603 firm-years versus our 820 for their sample period); and they use Compustat Global for data on shares outstanding while we rely on the Norwegian database *Børsprosjektet* (for Norwegian companies, the shares outstanding variable (CSHOC) in Compustat Global includes shares held in treasury, while we exclude treasury shares). Results from replicating Table 8 using a sample and methodology identical ti that of Ahern and Dittmar (2012) are available upon request.

[16] While not reported in Table 8, eliminating the five government-controlled firms alone eliminates the statistical significance of the slope coefficients in Panel B of Table 8 (reducing the coefficient to 1.6). Essentially, eliminating these firms reduces the cross-sectional variation in the second-stage regression, further underscoring the importance of measuring the first-stage initial value of the percentage female directors prior to 2002.

The vector $\mathbf{X}$ of control variables includes the total number of shareholder representatives on the board (*Board size*), firm age, the natural logarithm of the book value of total assets (*Total assets*), the ratio of book value of total debt to total assets (*Leverage*), the ownership fraction of the largest shareholder (*Large owner*), and a dummy variable indicating that the firm is OSE-listed (*Listed*). The regression also includes industry sector dummies. Moreover, since the regression relies on accounting data rather than stock returns, we expand the treatment group to include both listed and unlisted ASA.

Furthermore, we expand our sample to include a control group of "Large AS" firms, not regulated by the quota law. As before, our data source is the national corporate registry (Berner, Mjøs, and Olving, 2013). For unlisted firms we restrict the sample to stand-alone companies, excluding subsidiary firms that do not report consolidated accounts. As shown in Appendix Table 2, the population of active AS is 167,326 in an average sample year. Of these AS, 125,379 are stand-alone firms, of which 98,715 have the accounting information we require. Of these AS, we select the one percent largest firms by revenue each year, henceforth referred to as "Large AS". There are 987 Large AS in an average year, which is roughly twice the average number of ASA. The estimation of regression equation (9), however, excludes 142 Large AS that are registered as ASA at some point during the sample period.

Table 9 reports bi-annual firm and director characteristics for the samples of ASA and Large AS. As shown in Panel A, listed ASA are typically largest in both revenue and assets. Large AS (Panel C) are on average about half the size of listed ASA, but similar in size to (or slightly larger than) unlisted ASA (Panel B). Appendix Table 3 shows the industry sector distribution for the three subsamples, after assigning each firm-year to one of ten industry sectors. Listed ASA have the highest concentration in Offshore/Shipping and Telecom/IT/Tech, while unlisted ASA are concentrated in Finance and Other services industries. Large AS are particularly frequent in Construction and Wholesale/Retail.[17]

Table 10 shows the estimated regression coefficients for equation (9), using different sample periods and definitions of the post-quota period. The estimation period for column (1) is 1998-2013, excluding the quota-implementation years 2002-2008. Thus, this regression contrasts operating profitability before the first significant public discussion and after full implementation of the gender quota. The key variable of interest is $ASA \times Post2008$, which receives a statistically insignificant coefficient. This coefficient estimate fails to reject the null hypothesis of no impact of the gender quota on operating profitability.

[17] We thank Ida Kristine Skogvoll at Statistics Norway for assistance in creating a consistent definition of industry sectors across the sample period.

In column (2) of Table 10, we address the finding of Matsa and Miller (2013) of a significantly negative effect of the gender quota on ROA of 104 listed ASA. This column uses their sample period (2003-2009) and their definition of the post-implementation period ($Post2006$). Similar to Matsa and Miller (2013), we now find a negative coefficient on $ASA \times Post2006$, significant at the 5% level. Thus, it appears that ASA tend to under-perform (unregulated) Large AS in the 2007-2009 period. While our regression specification (9) differs somewhat from that of Matsa and Miller (2013), the coefficient estimate in column (2) supports their conclusion that the forced addition of female directors influenced board decisions negatively.

However, in Column (3) of Table 10 we maintain the Matsa-Miller post-quota definition while extending the sample period to 2013. From an econometric point of view, the longer sample period is advantageous because it reduces the relative impact of the financial crisis years on the estimated coefficients. Interestingly, this extension of the sample period drives the coefficient estimate for $ASA \times Post2006$ to a statistically insignificant value. A consistent interpretation is therefore that the negative coefficient on $ASA \times Post2006$ in column (2) is driven by abnormally low operating profits during the financial crisis. Thus, our larger treatment group (all Norwegian ASA) and longer sample period (through 2013) suggests a neutral impact of the gender quota on operating profitability.

The above analysis of short- and long-run performance suggests that the gender quota did *not* significantly affect the market value or the operating profitability of the typical ASA. In the remainder of the paper, we examine corporate actions such as legal conversions and board changes that provide further perspectives on this important conclusion.

4 Did the quota trigger ASA to AS conversions?

Norway introduced the ASA legal form in 1997 under its policy of complying with EU directives.[18] The total number of ASA peaked in 2001 and is continually changing as firms for various reasons switch between the ASA and AS forms. Major reasons include (but are not restricted to) a decision to go public (AS to ASA), acquisitions and liquidations (ASA to AS), and abandonment of plans to issue public equity (ASA to AS). In this section, we examine whether avoidance of the quota restriction is also an important driver of ASA to AS conversions, as claimed by both Ahern and Dittmar (2012) and Bøhren and Staubo

[18] While not an EU member, Norway, a highly integrated EU trading partner, adopts major EU directives.

(2014).

From an economic point of view, the key benefit of the ASA over the AS organizational form is that only the former corporate form may list on the OSE and sell equity to the public.[19] Thus, a conversion from *listed* ASA to AS entails the loss of both liquidity and a potentially important source of external finance.[20] A conversion from *unlisted* ASA to AS is much less costly, in particular if the firm's investment opportunity set has changed so as to make a public equity offering obsolete. The value of the option to issue public equity may have been reduced due to firm- and industry-specific investment factors as well as economy-wide developments. The latter include the burst of the technology "bubble" around 2000, which was followed by a 63% drop in the OSE internet technology index from 2001-2002, and the recent financial crisis.

If an *unlisted* ASA finds the public listing option sufficiently unattractive, it may be optimal to convert to AS for a number of reasons, primarily related to corporate governance. In particular, the conversion permits the firm to (i) combine the CEO and board chairmanship positions (illegal for an ASA), (ii) issue unlimited non-voting shares (the limit is 50% for ASA), (iii) avoid the more stringent rules on ownership and insider trading recording regulating ASA and, finally, (iv) avoid the gender quota.[21] The empirical challenge, next, is to determine whether this last factor is a driving force in the decision to convert to AS.

Column (1) of Table 11 shows the annual (year t) population of non-financial ASA, 2001-2007. We exclude financial firms from the analysis because these were required by law to be incorporated as ASA until 2007. Column (2) records the number of firms in year t exiting the ASA legal form in year $t+1$ *without* re-appearing as AS—due to forced and voluntary liquidation and merger. Column (3) records teh number of ASA in year t that convert to AS in year $t+1$ as subsidiaries after being acquired by a foreign and domestic firm (2002-2008). Since the events in Column (2) and (3) are hardly designed to avoid a gender quota, we net out these and concentrate the analysis on the sample in Column (4), where firms converted to from ASA to AS for all other reasons. These other reasons, which are generally not available in public data sources, may (hypothetically) include quota circumvention. The primary source for the information in Table 11 is the national *Brønnøysund* registry. In addition, for some of the conversions in

[19] An AS may, however, trade shares on the over-the-counter market and place equity issues privately.

[20] Other sources of finance—including private and public sale of debt—are equally available to ASA and AS (creditor protection laws do not distinguish ASA from AS).

[21] While listed ASA are required to adhere to the International Financial Reporting Standard (IFRS), unlisted ASA are not. Thus, relaxation of accounting standards is not a reason for an unlisted ASA to convert to AS.

Column (3), the acquirer is a foreign entity, which we identified using a manual search of other public data sources.

In Column (2), the reason for the exit is merger or forced or voluntary liquidation—all cases where the firm does not reappear as an AS. Column (3) shows the number of conversions from ASA to AS due to domestic or foreign acquisition and where the firm continues as subsidiary AS of the parent. Column (4) lists the number of firms converting from ASA to AS for all other reasons (generally unavailable in public data sources but may hypothetically include circumvention of the gender quota).

In a clarification of extant literature (Ahern and Dittmar, 2012; Bøhren and Staubo, 2014), column (4) of Panel A shows that not a single OSE-listed firm exited the ASA category for reasons other than bankruptcy, liquidation and acquisition.[22] Also, as shown both in Column (1) and in Panel A of Figure 2 above, the number of OSE-listed ASA increases each year from 2003 onwards, including over the two-year formal quota implementation period (2006-2007). We infer from this that, whatever the perceived cost of the gender quota, it was lower than the OSE listing benefit of the respective firms. If the gender quota helped trigger ASA to AS conversions, it did so for unlisted AS at best.

Column (4) of Panel B tracks conversions by unlisted ASA for reasons other than bankruptcy, liquidation and acquisition. Such conversions took place in 10% or 148 of the 1,417 firm-years over the period 2002-2008. While there is a slight increase in conversions during 2006 and 2007, the annual average number of conversions is unchanged at 21 both prior to (2002-2005) and during the formal quota compliance period (2006-2008). Without further testing, this suggests that conversions by unlisted ASA were driven fundamentally by incentives that existed in *all* sample years—not just during the formal quota compliance period. The relatively steady conversion rate does not rule out the possibility that some conversions occurred to circumvent the quota (recall the important quota news events already in 2002). However, it does suggest that a time series regression designed to test this particular incentive should allow for it to exist early in the sample period.

We estimate the following binomial logit model for the probability of converting from unlisted ASA to AS, using an unbalanced panel that starts in 2001 and ends in 2007:

$$Convert_{i,t} = \alpha + \gamma_1 Shortfall\ women_{i,t} + \gamma_2 \mathbf{X}_{i,t} + \epsilon_{i,t}, \tag{10}$$

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

where the dependent variable $Convert_{i,t}$ equals one in year t if the i'th unlisted ASA converts to AS in year $t+1$, and zero otherwise. All estimations include industry sector dummy variables, while Column (2) and (4) also include year fixed effects. This regression tests whether *Shortfall women* in year t drives the decision to convert in the following year, starting in year 2001. A firm is included in the sample up to the year prior to the year that it converts, and then dropped. The vector $\mathbf{X}_{i,t}$ of controls includes *Board size, Firm age, Total assets, ROA, Leverage, Largest owner*, as well as industry sector dummies. The unbalanced panel consists of the firm-years in Column (1) of Panel B in Table 11 minus the exits and conversions in Column (2) and (3), and requiring information on the firm characteristics $\mathbf{X}$. This yields a total of 821 firm-years: 480 firm-years for 148 converting firms in Column (4) and 441 firm-years for 127 firms in Column (1) that do not convert.[23]

The coefficient estimates for regression (10) are reported in Column (1) and (2) of Table 12, where Column (2) adds year fixed effects to the sector fixed effects in Column (1). Most important, the coefficient on *Shortfall women* is statistically insignificant, indicating that the unlisted ASA to AS conversions are unrelated to the gender quota. That is, there is no statistically significant difference in the value of *Shortfall women* between firms that convert (measured as of the year prior to the conversion) and those who do not (measured using all firm-years).

Columns (1) and (2) of Table 12 further shows that the likelihood of converting to AS increases with the ownership stake of the largest shareholder (*Largest owner*), and that it decreases with firm size (*Total assets*). The significance of these two variables is also interesting as several of the governance motives (i)-(iii) listed above likely play a greater role for smaller firms with relatively concentrated ownership. Of particular relevance is the the cost of a public equity offering, which are generally lower for larger firms with dispersed ownership (Eckbo, Masulis, and Norli, 2007). Consistent with this, and as suggested above, the positive impact of *Largest owner* and *Total assets* on the conversion probability may reflect a decision by the unlisted ASA to abandon a strategy of issuing public equity.

The above evidence is consistent with some of the results presented by Bøhren and Staubo (2014) but not with their main conclusion. Bøhren and Staubo (2014) base their main conclusion on a regression such as the one that we present in Column (3) of Table 12. In this column, when a firm converts to AS, the dependent variable is given a value of one for *all* prior years the firm is in the sample. As

[23] The inferences discussed below are basically unchanged if we also include the exits and conversions in Column (2) and (3) in the panel estimation.

reported by Bøhren and Staubo (2014) as well, this backfilling of the dependent variable produces a statistically significant positive coefficient estimate for *Shortfall woman*. However, the backfilling in Column (2) mechanically over-weighs *Shortfall women* in early pre-conversion years.[24] Consistent with this argument, Column (4) also shows that *Shortfall women* turns insignificant when we add year fixed effects to Column (3).

In sum, not a single listed ASA converted to AS for reasons other than merger, acquisition or liquidation. Moreover, the evidence in Table 12 fails to support the hypothesis that the gender quota caused even unlisted ASA, for which the opportunity cost of switching is particularly low, to convert to the AS legal form.

We next turn to large-sample evidence on actual board changes in both ASA and AS as a final check on the type of actions firms were observed to undertake around the quota implementation years.

5 The supply of female directors

The analysis so far shows insignificant short- and long-run firm performance effects of the gender quota (Table 3, 6, 7 and 10), a near-constant board size (Figure 1), and ASA to AS conversions that are statistically independent of the quota-induced shortfall of female directors (Table 12). Collectively, this evidence suggests that the pool of qualified female director was sufficiently deep for the quota law to be implemented without causing significant shareholder losses. As a final check on this conclusion, however, in this section we provide new large-sample evidence on changes in the female directorship pool that speaks more directly to the question of female director supply.

We consider the following three implications of the proposition that qualified female directors were (and perhaps continue to be) in short supply. The first is the emergence of a small group of relatively powerful, "busy" female directors. The second is an increase in female director turnover as many firms find their initial female directors to be under-qualified and take action to replace these with new females. The third implication is that director CEO and board experience should decline on average as the quota is implemented.

We examine these three implications using our sample of 402 listed ASA, 867 unlisted ASA, and 3,473 Large AS with the requisite board information over the period 1998-2013. These firms comprise

[24]Since firms that do not convert end up with a zero value for *Shortfall women*, the average value of *Shortfall women* is mechanically greater for converting than non-converting firms.

a total of 96,251 directorship-years covering 19,206 unique directors. In addition to a simple count of the number of directorships, we also construct for each director-year a measure of the director's network power using the network's PageRank score as defined in Appendix B. The PageRank score of an individual director increases with the "centrality" of the individual's network of board seats, which increases with the network size of the other board members in the individual's own network. We use the PageRank to describe the evolution of what we label the "gender power gap" over the sample period.

5.1 Directorship dispersion and network power

Figure 4 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 contains the total sample of all ASA directors (in both listed and unlisted ASA) and their directorships in ASA as well as in Large AS. As shown in Panel A, directorships in OSE-listed ASA are highly dispersed, and the dispersion of board seats is similar across male and female directors. Over the sample period, 85% of all directors sit on a single board and 11% sit on two boards.

Ferris, Jagannathan, and Pritchard (2003) and Field, Lowry, and Mkrtchyan (2013) define a director holding three or more board seats as "busy". With this definition, only 4% of the directors of OSE-listed firms are busy in an average year. The fraction of busy directors increases only slightly for female directors—from 3% in 1998-2005 to 6% in 2008-2013—while it decreases from 4% to 2% male directors. As expected, including unlisted ASA and Large AS (Panel B) produces higher average fractions of busy ASA directors. However, the *change* in this fraction—the key statistic for our purpose—remains small. For females, the fraction increases from 7% pre-quota to 12% post-quota, while the fraction increases from 12% to 15% for males. In sum, there is little evidence that the gender quota has increased directorship concentration among directors of listed ASA, regardless of their gender.[25]

Turning to the PageRank director power score, Table 13 lists the annual distribution of the male and female ASA directors' scores and the power gap, 1998-2013. The power gap is defined as one minus the ratio of female to male director power. As shown, the male and female director power are increasing over

[25] Interestingly, notwithstanding the generally much larger firm sizes of US listed companies, Cashman, Gillan, and Jun (2012) classify as much as 25% of the directors of S&P1500 companies as 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 sample period, with a slight decline in the median value of the power gap. At the end of our sample period, in 2013, both the mean and median values of the power gap is 7%, suggesting a relatively small but positive male power gap five years after gender quota implementation. Overall, both a direct count of board seats and the PageRank power score indicate a wide male and female board seat dispersion both before and after the quota.

5.2 Director turnover

Figure 5 plots the annual fraction of male and female ASA directors that leave the board next year. The turnover rate is similar for men and women, averaging 20%-25% over the entire sample period (1998-2013). Recall from Figure 1 that the average OSE-listed ASA has five shareholder-elected directors throughout the sample period, of which 1.2 were female at the end of 2005. Thus, with a typical annual turnover rate of 25%, this firm was in a position to comply with the quota requirement over the two-year compliance period well within its normal turnover rate. This likely explains why Figure 5 show no (univariate) evidence of an increase in female director turnover in the years surrounding mandatory quota compliance.

We use the following multivariate regression to test for a quota-induced increase in female director turnover more formally:

$$\textit{Board turnover}_{i,t} = \alpha + \gamma_1 \text{ASA x Post2008}_{i,t} + \gamma_2 \text{Post2008}_{i,t} + \gamma_3 \text{ASA}_{i,t} + \gamma_4 \mathbf{x}_{i,t} + \epsilon_{i,t}, \tag{11}$$

where *Board turnover* is the fraction of directors of firm i at year-end t that have left at year-end t+1. The explanatory variables in $\mathbf{x}$ include the dummy *Implementation* indicating years 2001-2007 for turnover in 2002-2008 (the full quota implementation period), average *Female board power* computed using Page Rank, average *Board power*, dummy variables indicating *Female CEO* and *Female chair*, and *ROA* (ratio of EBIT to book value of assets). The remaining firm characteristics and time-period dummies in $\mathbf{x}$ are as defined earlier in Table 10. The sample consists of 855 ASA and 3,060 Large AS, excluding financial firms and Large AS registered as ASA at some point during the sample period, for a total sample of 13,621 firm-years in 1998-2013.

Table 14 shows the coefficient estimates. The dependent variable in Column (1) is total board turnover, while it is male and female director turnover, respectively, in Column (2) and (3). The coefficient for

$ASA \times Implementation$ is significant and positive in Column (1) and (2), suggesting that male director turnover increased during the implementation of the quota, perhaps to give room for the incoming female directors. However, neither $ASA \times Post2008$ nor $ASA \times Implementation$ are significant in Column (3), indicating that there was no abnormal turnover of female ASA directors as the quota was implemented or afterwards.

As expected, *Board power* enters with a significantly negative coefficient in all three regressions, likely because powerful boards tend to have above-average-quality directors and thus lower turnover. Moreover, having controlled for overall board power, *Female board power* generates an insignificant coefficient estimate, suggesting that the impact of director network power on turnover is similar across men and women. Female director turnover is, however, lower in firms with a female CEO. One reason might be that female CEOs are better at spotting and hiring high-quality female directors. It is also possible that (the rare) organization with a female CEO tends to exhibit greater acceptance of female directors regardless of their quality.

Several of the remaining regression coefficients are also significant, indicating that director turnover is increasing in board size and tends to be higher for listed ASA and for younger firms. Also, turnover increases with the ownership stake of the largest shareholder, suggesting greater shareholder monitoring of the board in firms with concentrated ownership. As expected, directors are also more likely to leave firms with poor operating performance (ROA) and high leverage (indicating financial distress).

Overall, there is no evidence in Table 14 of increased female director turnover as a result of gender quota implementation. We next examine whether the quota law has reduced overall board quality, as reflected by the skills and experience of individual directors.

5.3 Director experience

If firms are forced to hire under-qualified female directors, we should see a decline in measures of overall board experience. Figure 6 plots the annual average director CEO experience (Panel A) and cumulative directorship-years for ASA directors, 2001-2013. The figure displays these measures for the total board as well as for male and female directors separately. In Panel A, CEO experience is defined as being a current outside CEO or past CEO (since 1998) of an ASA or a Large AS. As expected, the fraction of ASA directors with CEO experience is much lower among women than men. Moreover, the level of CEO experience of female directors remains relatively low throughout the sample period, at about 10%.

Interestingly, notwithstanding the low CEO experience of female directors, the average level of board CEO experience in Panel A does *not* decline as the quota is implemented. This occurs because the low female CEO experience is offset by a steadily increasing fraction of male directorships with CEO experience, from a low of 25% in 2001 to a high of 40% in 2013. In other words, facing forced election of female directors with relatively low CEO experience caused shareholders to replace the *least* experienced male directors to the point where the overall board experience is unchanged on average. This director replacement strategy supports our finding above of a negligible overall valuation effect of the gender quota.

Panel B of Figure 6 plots our second measure of individual director experience for ASA directors: the number of directorship-years in year t cumulated over the previous three years (years $t-3$ through $t-1$). Consistent with Figure 4 above, there is a near-complete convergence of male and female cumulative directorship in the years following quota implementation. This complements the evidence on director CEO experience in that the overall board-level experience does not appear to have suffered for the average OSE-listed firm implementing the gender quota requirement.

Finally, returning to Table 9 above, there is evidence that the gender quota has had spill-over effects into the C-suite (senior management) as well as on unregulated firms. The table summarizes the evolution of the fraction of female directors, board chairs, and CEOs for both ASA and Large AS. First, the fraction of female directors in Large AS has increased somewhat, from 7% in 1998 to 13% in 2013. Second, the fraction of female board chairs has also increased, from about 1% in 1998 to 9% of listed ASA and 15% of unlisted ASA in 2013. Third, 5% of listed ASA and 10% of unlisted ASA have female CEOs in 2013, up from 3% and 2%, respectively, in 1998. While a deeper analysis of the likely source of these changes goes beyond this paper, it is certainly possible that the quota law has impacted the gender-balance of firms beyond those directly constrained by the law, as the legislators originally intended.[26]

6 Conclusion

The past fifteen years have seen a political trend in western liberal democracies towards increased gender-balancing of corporate boards, voluntarily or through legislation. In this paper, we present new empirical

[26] An early government white-paper, *Odelstingsproposisjon* 97, 2002-2003 (submitted to the Norwegian Parliament by the Ministry preparing for the gender quota law), cites a broad objective of increasing female power in business. See also Bertrand, Black, Jensen, and Lleras-Muney (2014) for an empirical examination of wider societal effects of the quota.

tests of whether gender balancing affects firm value. Our experimental setting is the 2005 Norwegian gender quota law. This law requires each gender to hold about 40% of the shareholder-elected board seats of ASA (public limited liability) companies, and firms were given until the end of 2007 to comply. Several other European countries have since passed their own gender quotas.

By being the first of its kind, and because it came with a severe sanction for non-compliance (forced liquidation), Norway's quota represents an ideal laboratory for identifying valuation effects of exogenous changes in board composition and independence. The main question addressed by our empirical analysis is whether forced gender balancing, which constrains the free choice of board composition and lowers average director experience in the short run, produces a statistically and economically significant reduction in the market value of firms listed on the OSE.

One immediate effect of the quota was to increase director independence since, as we show, the incoming female directors score lower on director network power than do male directors. While the governance literature tends to view director independence as positive for shareholders, the question for this paper is whether this independence is gained at the expense of director qualifications and experience. Contrary to extant research, our robust large-sample analysis fails to reject the hypothesis of a value-neutral effect of the quota law. This conclusion emanates from a comprehensive event-study of quota-related news announcements, as well as from long-run analyses of stock and accounting performance.

Our finding of a value-neutral effect is also consistent with other aspects of the behavior of firms surrounding the quota implementation period. For example, firms neglected to expand board size, which would have been a relatively cheap and effective option to give room for new females without having to sacrifice valued male directors or board control. Moreover, not a single OSE-listed firm decided to convert to the AS legal form—which exempts the firm from the quota restriction—without being acquired. Also, while we observe several ASA to AS conversions by *unlisted* ASA over the sample period, our analysis indicates that these conversions were unrelated to the quota-induced female director shortfall.

Also interesting, we provide new, large-sample evidence on the evolution of board seat dispersion and director turnover and experience. This evidence sheds light on whether qualified female directors were in short supply when the quota was implemented. With qualified females in short supply, one would expect board seats to become concentrated among a few but busy female directors. Moreover, one would expect director turnover among female directors to be relatively high as some boards attempt to replace under-qualified female directors *ex post*.

We find instead that director seat dispersion is wide for both men and women, and has remained largely unaffected by the quota. Moreover, there is no evidence that female director turnover on average increases post-quota, consistent with a sufficiently deep pool of qualified female director candidates. Also interesting, while incoming female directors have far less CEO experience than their male counterparts, we show that boards nevertheless have managed to maintain a relatively constant overall CEO experience.

In sum, our findings of (i) insignificant abnormal stock and accounting performance (both short- and long-run), (ii) lack of board size expansion, (iii) ASA to AS conversion that are statistically unrelated to the female shortfall generated by the quota, (iv) wide male and female board seat dispersion both before and after the quota, (v) normal post-quota female director turnover rates, and (6) relatively constant overall director CEO experience for ASA boards, all point to a value-neutral effect of the Norwegian gender quota for OSE-listed firms. Shareholder-borne costs associated with the forced gender-balancing appear to have been modest.

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Figure 1
Average board size and fraction of female directors in listed ASA, 1998-2013

The figure shows the annual average board size, fraction of female directors, and number of shareholder-elected and female directors, for the population of 402 unique ASA firms (with board information) listed on the Oslo Stock Exchange (OSE), 1998-2013. The board data is from *Brønnøysund Register Centre*. The two vertical lines indicate the quota being signed into law (year-end 2005) and mandated compliance (year-end 2007).

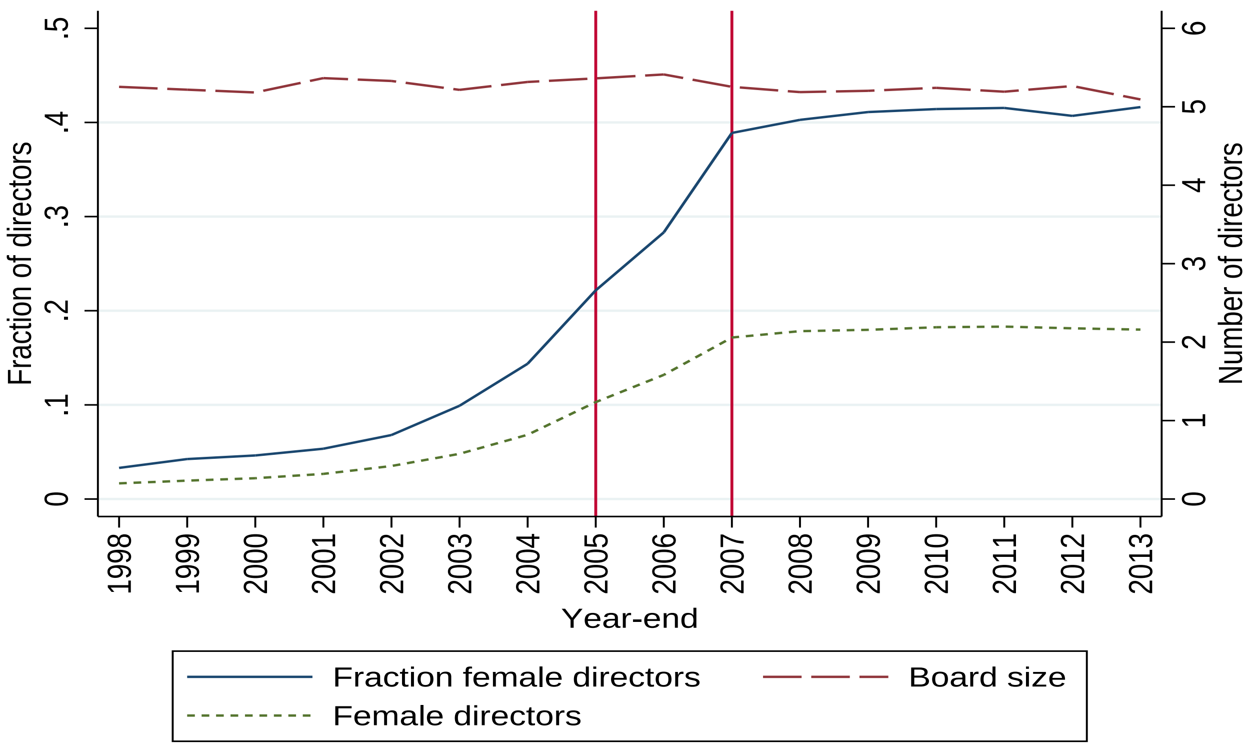
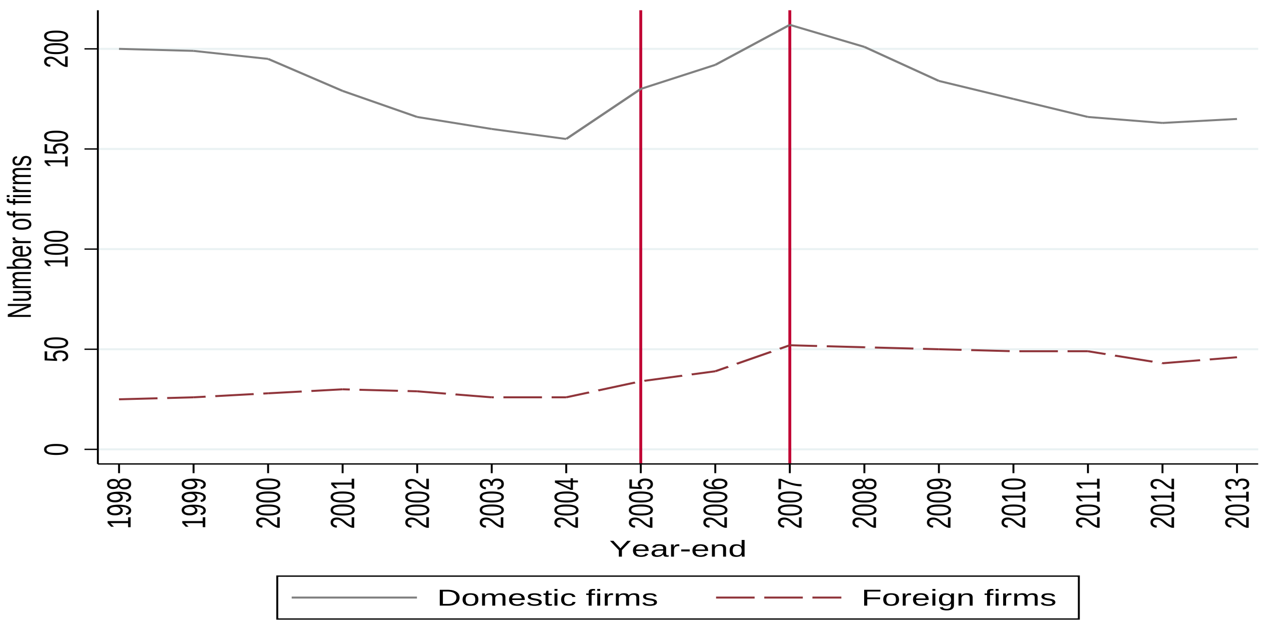


Figure 2
Number of firms listed on OSE and in the event-study portfolios, 1998-2013

Panel A shows the annual population of domestic and foreign firms listed on Oslo Stock Exchange (OSE, including *Oslo Axess* from 2007) between year-ends 1998 and 2013. The two vertical lines mark the quota being signed into law (year-end 2005) and mandated quota compliance (year-end 2007). Panel B plots the daily number of sample firms in the two portfolios of domestic and foreign OSE-listed firms used to estimate quota-induced abnormal stock returns.

A: Number of domestic and foreign OSE-listed firms



B: Number of domestic and foreign firms in the event-study portfolios

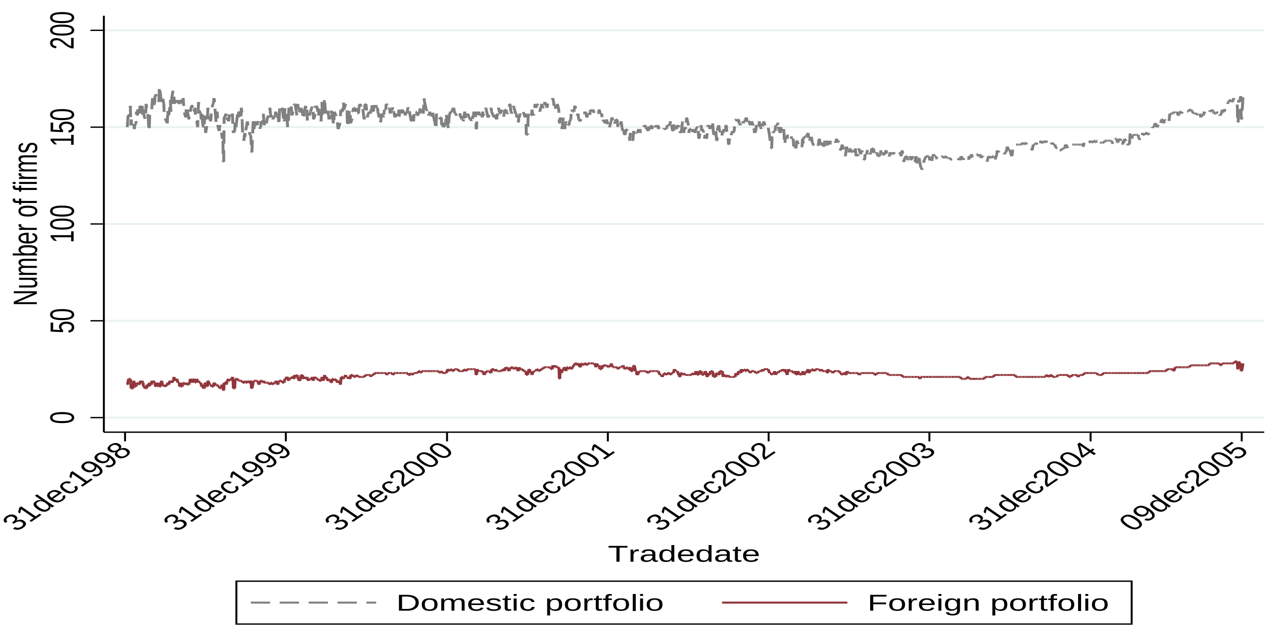


Figure 3
Monthly cumulative raw return for low-shortfall and high-shortfall portfolios

The figure shows the monthly cumulative raw return for portfolios of low-shortfall and high-shortfall ASA firms listed on the Oslo Stock Exchange, from February 2002 (start of significant quota legislation news) to April 2008 (full quota compliance). A low-shortfall firm has at least one female director in 2001, while a high-shortfall firm has zero female directors in 2001. The monthly average number of firms in the low-shortfall and high-shortfall portfolios is 32 and 98, respectively. The figure also shows the monthly cumulative return on the MSCI world stock market index converted to NOK.

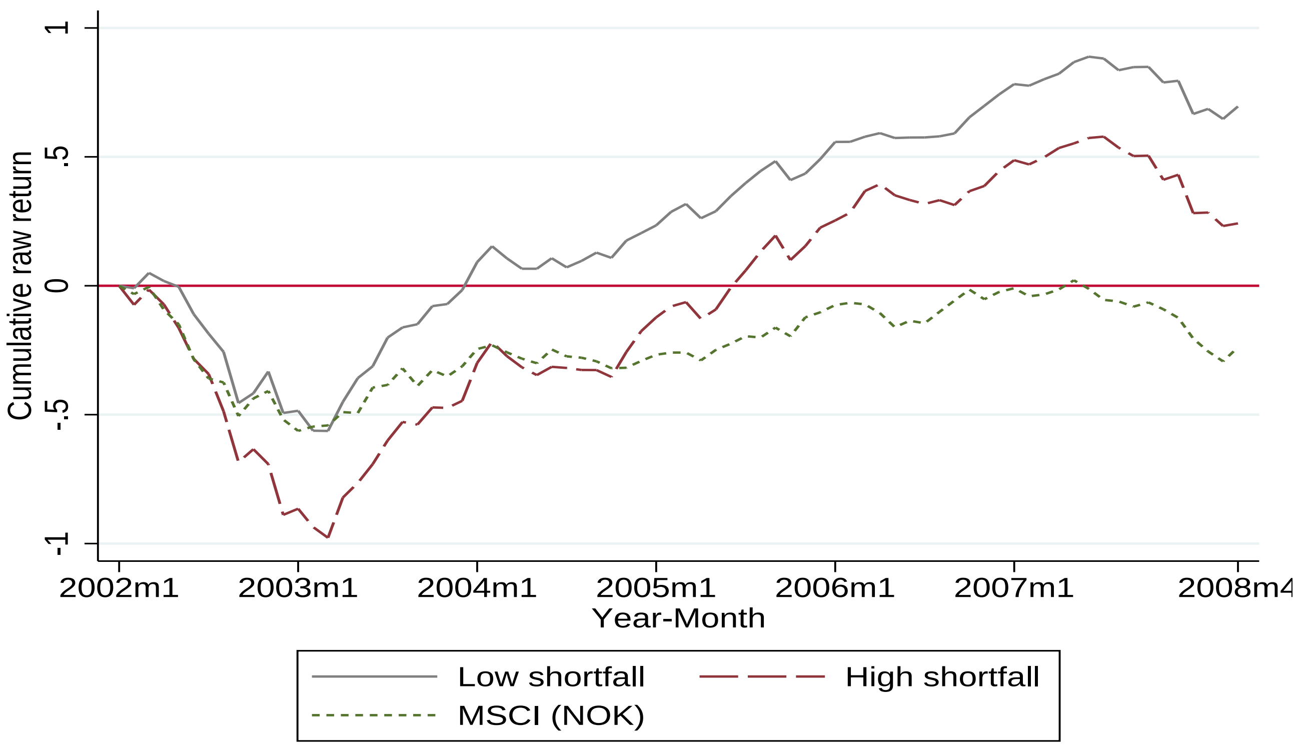
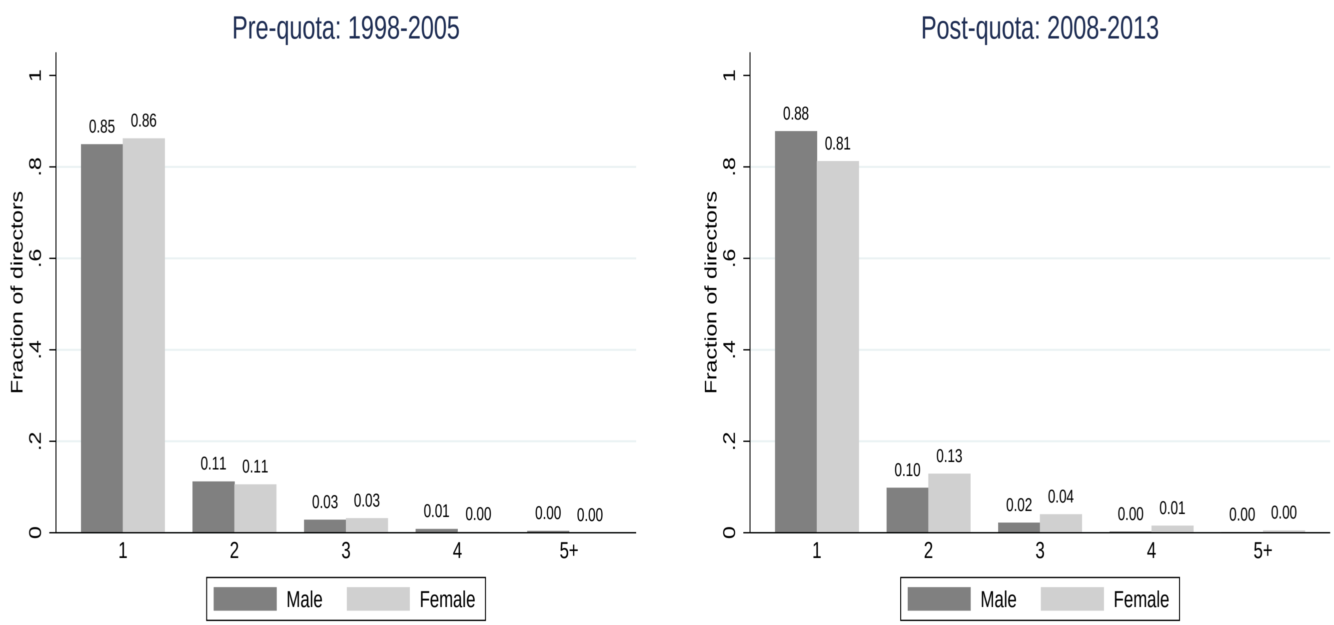


Figure 4
Distribution of ASA directors by their number of board seats, 1998-2013

The figure plots the frequency distribution of male and female ASA directors by the number of boards they serve on. Panel A counts directorships in the population of listed ASA firms, while Panel B counts directorships in the population of ASA (listed and unlisted) and Large AS (the top 1% stand-alone AS by revenue). Five and more directorships are reported under 5+. Each panel shows the average distribution for the periods 1998-2005 (before the quota is formally mandated) and 2008-2013 (after full compliance is achieved). The sample is 96,251 directorship-years for 19,206 unique directors in 402 listed ASA, 867 unlisted ASA, and 3,473 Large AS (with board information).

Panel A: Distribution of listed ASA directors by their number of directorships in listed ASA



Panel B: Distribution of all ASA directors by their number of directorships in ASA and Large AS

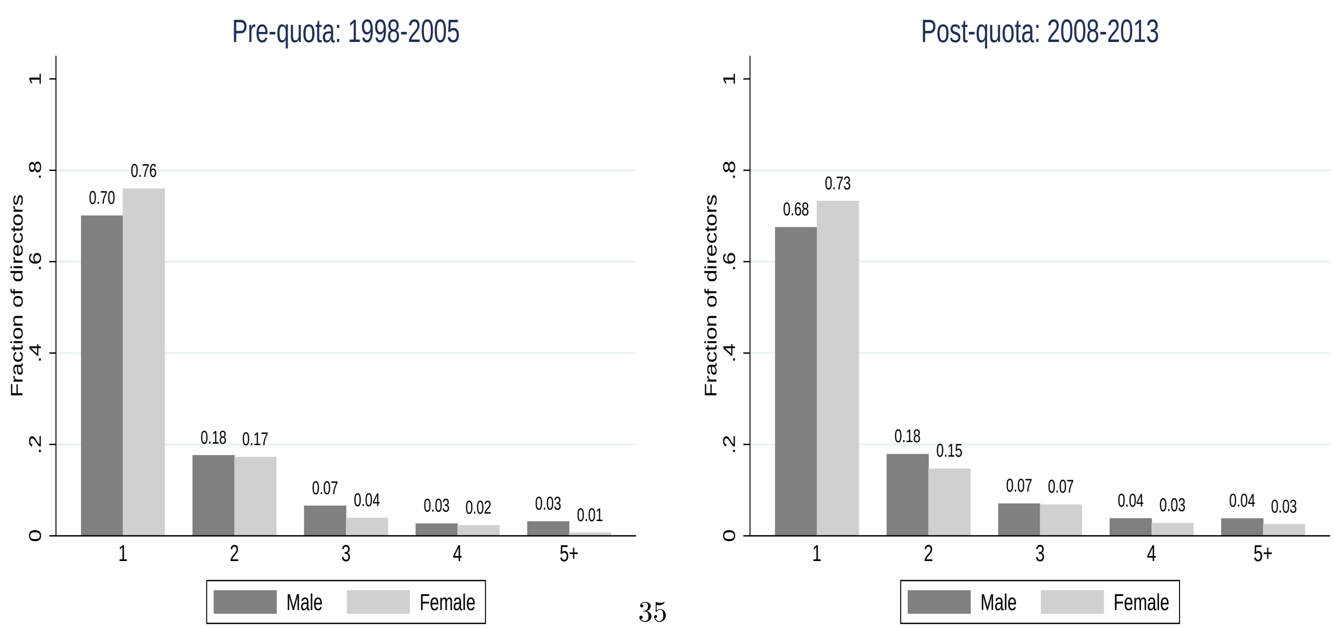


Figure 5
Turnover rate of ASA directors, 1998-2013

The figure shows the annual turnover rate for the population of male and female ASA directors, 1998-2013. Turnover is defined as the fraction of the directors in firm i at year-end t that are no longer on the board of firm i at year-end $t+1$. The two vertical lines mark the quota being signed into law (year-end 2005) and mandated compliance (year-end 2007), The sample is 96,251 directorship-years for 19,206 unique directors in 402 listed ASA, 867 unlisted ASA, and 3,473 Large AS (with board information).

Figure 6
CEO and director experience of ASA directors, 1998-2013

Panel A shows the annual fraction of ASA boards, and of male and female ASA directorships, with CEO experience, defined as a current outside CEO or past CEO (since 1998) at an ASA or Large AS. Panel B shows the annual average directorship experience for boards, and for male and female ASA directorships, 2001-2013. Directorship experience at end of year *t* is defined as the cumulative number of directorship-years from year-end *t-3* to year-end *t-1* in ASA and Large AS. The two vertical lines indicate the quota being signed into law (year-end 2005) and mandated compliance (year-end 2007).The sample is 96,251 directorship-years for 19,206 unique directors in 402 listed ASA, 867 unlisted ASA, and 3,473 Large AS (with board information).

Panel A: Director CEO experience for ASA boards

Panel B: Average cumulative directorship-years for ASA directors

Table 1
Fraction of female directors required by Norway's board gender quota

The table shows how the minimum number and fraction of women (and men) required by Norway's 2005 gender quota varies with board size, here the number of shareholder-elected directors. The quota regulates the shareholder-elected directors of ASA firms, about half of which are listed on the Oslo Stock Exchange during the sample period 1998-2013. The quota was mandated in 2005 and required full compliance by year-end 2007.

Board size (shareholder-elected directors)	Minimum number of female (and male) directors	Minimum fraction of female (and male) 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
Key news events leading up to the passage of Norway's board gender quota

This table describes key news events leading up to the signing of the board gender quota into Norwegian corporate law in 2005. The panels divide this process into four stages. Panel A shows early discussions of a corporate board quota in gender equality law. In Panel B, the discussion switch to a corporate board quota in corporate law. At the end of Panel B (in July 2001), however, the cabinet faces fall elections (which eventually led to a change of government) and lacks time for the necessary legislative process. Panel C contains, for the first time, major developments toward a parliamentary majority in favor of the quota. Finally, in Panel D, the quota was signed into law on December 9, 2005. Newspaper sources are shown in parentheses.

A: Early discussions of a corporate board gender quota in gender equality law

(1) August 12, 1999: Cabinet press conference on proposed amendments to gender equality law extending a 40% gender quota for public committees to corporate boards (*Dagens Næringsliv*).

(2) October 15, 1999: Cabinet releases details of the proposed amendments to gender equality law for public consultation.

B: Early steps toward a board gender quota in corporate law

(3) March 8, 2001: Cabinet presents an amendment to gender equality law that will be sent to Parliament for a vote, but which does *not* include a board gender quota (*Verdens Gang*).

(4) July 2, 2001: With two months left to elections, the center-left Cabinet releases for public consultation a proposed board gender quota for ASA firms to be amended to corporate law.

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

(5) February 22, 2002: The newly elected (center-right) Minister of Trade and Industry surprisingly supports a gender quota in a newspaper interview (*Verdens Gang*).

(6) February 23, 2002: The next day, the same Minister retracts his support in a newspaper interview (*Dagens Næringsliv*).

One week later, the parliamentary members of his conservative party rejects a quota, causing much public debate which, surprisingly and after negotiations with the other parties in the government, culminates in a decision to introduce a quota law.

(7) March 8, 2002: The center-right Cabinet proposes a board gender quota to be signed into law in 2005 absent voluntary compliance. The Cabinet promises compliance by government-owned firms within one year (*Dagens Næringsliv*).

(8) June 13, 2003: Cabinet's gender quota proposal submitted to Parliament.

(9) November 28, 2003: The evening before, Parliament votes in favor of the gender quota amendment to Corporate Law. To become law, the amendment must be signed by the Cabinet which, due to the voluntary compliance option, will occur in September 2005 at the earliest.

D: Signing of the board gender quota into corporate law

(10) December 1, 2005: Absent sufficient change in board gender diversity, the Prime Minister states that the Cabinet will sign the gender quota amendment into law with a monetary sanction. (*Verdens Gang*).

(11) December 9, 2005: Cabinet signs the quota amendment into law. The ultimate sanction for non-compliance is forced liquidation—as for any breach of Corporate Law. Existing ASA firms have two years to comply.

Table 3
Abnormal stock returns to portfolios of OSE-listed firms on key quota event dates

Panel A reports abnormal stock returns for portfolios of domestic and foreign OSE-listed firms over the two-day event window k, $2 \times AR_k$, estimated using the following two different return-generating processes:

$$r^e_{k,t} = \alpha_k + AR_k d_{k,t} + \varepsilon_{k,t}, \quad \text{and} \quad r^e_{k,t} = \alpha_k + AR_k d_{k,t} + \beta_{k1} W^e_t + \beta_{k2} W^e_{t-1} + \varepsilon_{k,t},$$

where $r^e_{k,t}$ is the daily equal-weighted return in excess of the one-day Norwegian interbank offer rate (NIBOR), $d_{k,t}$ is a dummy for the event window (-1,0), and W^e is the daily return on the MSCI stock market world index (converted to NOK with the daily exchange rate) in excess of NIBOR. The eleven event days $k = 1, ..., 11$ are defined in Table 2. The event window for event (6) is (0,1) to avoid overlap with event (5). The abnormal return to each event is estimated separately, using daily portfolio returns over days -251 through day 0. The estimation of the first event starts on December 31, 1998. To be included in the event portfolio, a firm must have a minimum of 100 return observations in the estimation period, and it must have return observation in both days of the two-day event window. For each of the eleven event portfolios, we dummy out earlier event dates in the estimation period. In Panel B, subscript k is redefined to denote one of four legislative subperiods, and the portfolio abnormal return is estimated by letting the event dummy $d_{k,t}$ take a value of one in all of the two-day event windows covered by the k'th legislative subperiod. Thus, for each legislative period, the cumulative abnormal return reported in Panel B is $CAR_k \equiv K_k \times AR_k$, where K_k is twice the number of events in the k'th subperiod ($k = 1, .., 4$). The legislative subperiod estimation starts 251 trading days prior to the first event in the subperiod and ends on day 0 of the last event. If the estimation period includes events from prior legislative subperiods, then these prior event windows (-1,0) are dummied out. The sample comprises 339 unique OSE-listed firms, of which 290 are domestic and 49 are foreign. Robust standard errors (White estimator) in parenthesis. Significance levels are *** 1%, ** 5%, * 10%.

	$r^e_t = \alpha + AR_k d_{k,t} + \varepsilon_t$			$r^e_t = \alpha + AR_k d_{k,t} + \beta_1 W^e_t + \beta_2 W^e_{t-1} + \varepsilon_t$		
	Domestic firms (1)	Foreign firms (2)	Domestic –Foreign (3)	Domestic firms (4)	Foreign firms (5)	Domestic –Foreign (6)
A: Two-day abnormal returns ($2 \times AR_k$) for key news events in Table 2						
(1) August 12, 1999	0.006	-0.001	0.008	0.003	-0.005	0.007
	(0.002)	(0.002)	(0.003)	(0.003)	(0.002)	(0.003)
(2) October 15, 1999	-0.029	-0.032	0.003	-0.021	-0.024	0.003
	(0.010)	(0.012)	(0.002)	(0.008)	(0.010)	(0.003)
(3) March 8, 2001	-0.001	-0.002	0.000	-0.003	-0.005	0.002
	(0.002)	(0.003)	(0.005)	(0.002)	(0.004)	(0.005)
(4) July 2, 2001	0.011	0.014**	-0.003	-0.003	-0.013***	0.010
	(0.006)	(0.004)	(0.009)	(0.007)	(0.002)	(0.009)
(5) February 22, 2002	-0.008	-0.006	-0.002	-0.009**	-0.007**	-0.002
	(0.003)	(0.002)	(0.001)	(0.002)	(0.002)	(0.001)
(6) February 23, 2002	0.006	0.021	-0.015**	-0.006	0.003	-0.008
	(0.003)	(0.007)	(0.003)	(0.005)	(0.008)	(0.003)
(7) March 8, 2002	0.031***	0.041***	-0.010	0.028***	0.035***	-0.007
	(0.002)	(0.003)	(0.006)	(0.002)	(0.005)	(0.007)
(8) June 13, 2003	-0.007***	0.005	-0.011	-0.006	0.004	-0.010
	(0.001)	(0.004)	(0.005)	(0.003)	(0.007)	(0.005)
(9) November 28, 2003	-0.003	0.007	-0.010	0.003	0.012	-0.010
	(0.002)	(0.009)	(0.008)	(0.003)	(0.010)	(0.008)
(10) December 1, 2005	0.010***	-0.002	0.012**	0.007***	-0.006	0.013**
	(0.001)	(0.003)	(0.003)	(0.001)	(0.002)	(0.003)
(11) December 9, 2005	0.005	0.013***	-0.008***	0.005	0.013***	-0.008***
	(0.002)	(0.001)	(0.001)	(0.003)	(0.002)	(0.001)
B: Cumulative abnormal returns (CAR) for each legislative subperiod in Table 2						
Subperiod A (events 1+2)	-0.022	-0.032	0.010	-0.017	-0.028	0.011
	(0.007)	(0.007)	(0.002)	(0.005)	(0.006)	(0.002)
Subperiod B (events 3+4)	0.011	0.013	-0.002	-0.005	-0.015*	0.010
	(0.003)	(0.003)	(0.005)	(0.004)	(0.002)	(0.005)
Subperiod C (events 5+6+7+8+9)	0.016	0.065*	-0.049**	0.009	0.052	-0.043*
	(0.002)	(0.004)	(0.002)	(0.003)	(0.004)	(0.002)
Subperiod D (events 10+11)	0.015**	0.011	0.004	0.012**	0.007	0.005
	(0.001)	(0.003)	(0.003)	(0.001)	(0.003)	(0.003)

Table 4
Abnormal stock returns to OSE-listed firms using a multi-factor model

The table reports abnormal stock returns for portfolios of domestic and foreign OSE-listed firms over the two-day event window k, $2 \times AR_k$, estimated using the following return-generating process:

$$r^e_{k,t} = \alpha_k + AR_k d_{k,t} + \beta_{k1} W^e_t + \beta_{k2} W^e_{t-1} + \beta_{k3} HML_t + \beta_{k4} SMB_t + \beta_{k5} MOM_t + \varepsilon_{k,t}, \quad t = -251, .., 0,$$

where $r^e_{k,t}$ is the daily equal-weighted return in excess of the one-day Norwegian interbank offer rate (NIBOR), $d_{k,t}$ is a dummy for the event window (-1,0), and W^e is the daily return on the MSCI stock market world index (converted to NOK with the daily exchange rate) in excess of NIBOR. The factors HML and SMB are the value and size factors (Fama and French, 1992) and MOM is the momentum factor (Carhart, 1997), calculated using the population of OSE stocks and made available by Bernt A. Odegaard on his web site, http://finance.bi.no/∼bernt/. The eleven event days $k = 1, ..., 11$ are defined in Table 2. The event window for event (6) is (0,1) to avoid overlap with event (5). The abnormal return to each event is estimated separately, using daily portfolio returns over days -251 through day 0. The estimation of the first event starts on December 31, 1998. To be included in the event portfolio, a firm must have a minimum of 100 return observations in the estimation period, and it must have return observation in both days of the two-day event window. For each of the eleven event portfolios, we dummy out earlier event dates in the estimation period. In Panel B, subscript k is redefined to denote one of four legislative subperiods, and the portfolio abnormal return is estimated by letting the event dummy $d_{k,t}$ take a value of one in all of the two-day event windows covered by the k'th legislative subperiod. Thus, for each legislative period, the cumulative abnormal return reported in Panel B is $CAR_k \equiv K_k \times AR_k$, where K_k is twice the number of events in the k'th subperiod ($k = 1, .., 4$). The legislative subperiod estimation starts 251 trading days prior to the first event in the subperiod and ends on day 0 of the last event. If the estimation period includes events from prior legislative subperiods, then these prior event windows (-1,0) are dummied out. The sample comprises 339 unique OSE-listed firms, of which 290 are domestic and 49 are foreign. Robust standard errors (White estimator) in parenthesis. Significance levels are *** 1%, ** 5%, * 10%.

	Domestic firms (1)	Foreign firms (2)	Domestic - Foreign (3)
A: Two-day abnormal returns ($2 \times AR_k$) for key news events in Table 2			
August 12, 1999	0.004	-0.001	0.004
	(0.001)	(0.004)	(0.005)
October 15, 1999	-0.023	-0.030	0.006
	(0.009)	(0.011)	(0.004)
March 8, 2001	-0.000	-0.003	0.003
	(0.003)	(0.003)	(0.006)
July 2, 2001	-0.004	-0.013**	0.009
	(0.006)	(0.003)	(0.009)
February 22, 2002	-0.014***	-0.013***	-0.001
	(0.001)	(0.001)	(0.001)
February 25, 2002	0.003	0.010	-0.006
	(0.003)	(0.005)	(0.003)
March 8, 2002	0.028***	0.032***	-0.004
	(0.001)	(0.006)	(0.006)
June 13, 2003	-0.001	0.008	-0.008
	(0.005)	(0.008)	(0.004)
November 28, 2003	0.000	0.011	-0.010
	(0.004)	(0.011)	(0.008)
December 1, 2005	-0.003	-0.023	0.021***
	(0.005)	(0.009)	(0.004)
December 9, 2005	0.008***	0.020***	-0.012
	(0.001)	(0.003)	(0.004)
B: Cumulative abnormal returns (CAR) for each legislative subperiod in Table 2			
Panel A (events 1+2)	-0.019	-0.032	0.013
	(0.005)	(0.007)	(0.003)
Panel B (events 3+4)	-0.004	-0.015	0.011
	(0.003)	(0.003)	(0.005)
Panel C (events 5+6+7+8+9)	0.005	0.047	-0.042*
	(0.002)	(0.004)	(0.002)
Panel D (events 10+11)	0.006	-0.002	0.008
	(0.003)	(0.007)	(0.005)

Table 5
Cross-sectional determinants of abnormal returns in legislative subperiods

This table reports coefficient estimates in cross-sectional regressions with firm i's abnormal return cumulated over the events in legislative subperiod k, $CAR_{i,k}$, as dependent variable ($k = 1, .., 4$). $CAR_{i,k}$ is estimated using Eq. (2) as return generating process, as in column (4) in Panel B of Table 3. The estimation period for $CAR_{i,k}$ starts 251 days before the first event in the subperiod and ends with the last event. If the estimation period includes events from prior legislative subperiods, then these prior event windows (-1,0) are dummied out. To be included, a firm must have return observations on all event days and $\geq$100 return observations in the estimation period. The OLS regressions is:

$$CAR_{i,k} = \alpha_k + \beta_k \mathbf{X_{i,k}} + u_{i,k}, \qquad k = 1, ...4,$$

where the vector $\mathbf{X}$ contains six firm characteristics: *Shortfall women* is the quota-induced shortfall of female directors, computed as the maximum of zero and the difference between the quota requirement (column (3) of Table 1) and the actual fraction of female directors. *Codetermination* is a dummy indicating that (quota-induced female directors+employee directors)/board size) $\geq$ 50%. *Risk* is the firm's daily stock return volatility in the prior year. *Firm size* is the natural logarithm of market value of equity. *Government control* is a dummy indicating that the government is the largest owner. *Largest owner* is the percent equity-ownership of the largest shareholder. All variables are from the year-end prior to the first event in the respective legislative subperiod. All regressions contain industry sector dummies. Robust standard errors (White estimator) are reported in parenthesis. Stars indicate significance levels: *** 1%, ** 5%, * 10%.

	Period A: First public hearing (Events 1-2) (1)	Period B: Second public hearing (Events 3-4) (2)	Period C: Law proposal (Event 5-9) (3)	Period D: Quota mandated (Event 10-11) (4)
Shortfall women	0.008	-0.047	-0.079	-0.003
	(0.066)	(0.072)	(0.095)	(0.026)
Codetermination	0.007	0.022	-0.017	0.004
	(0.018)	(0.019)	(0.027)	(0.008)
Risk	0.021	-0.945	-0.918	-0.828
	(0.492)	(0.606)	(0.738)	(0.538)
Government control	0.030	-0.008	-0.034	-0.020
	(0.019)	(0.032)	(0.049)	(0.016)
Largest owner	-0.001	0.031	-0.038	-0.009
	(0.022)	(0.053)	(0.056)	(0.026)
Firm size	-0.002	-0.018***	0.007	0.003
	(0.005)	(0.005)	(0.010)	(0.003)
Sector dummies	Yes	Yes	Yes	Yes
R^2	0.18	0.13	0.15	0.18
Number of firms	113	127	103	125

Table 6
Calendar time portfolio returns for low-shortfall and high-shortfall portfolios

The table reports monthly abnormal stock returns for portfolios of low-shortfall and high-shortfall ASA firms listed on the Oslo Stock Exchange, from February 2002 (start of quota legislative process) to April 2008 (full quota compliance). A low-shortfall firm has at least one female director in 2001, while a high-shortfall firm has zero female directors in 2001. The monthly average number of firms in the low-shortfall and high-shortfall portfolios is 32 and 98, respectively. In columns (1)-(3), the abnormal stock return is estimated using the following return-generating processes:

$$r_t^e = \alpha + \beta_1 W_t^e + \beta_2 W_{t-1}^e + \varepsilon_t,$$

where r_t^e is the monthly equal-weighted portfolio return on day t in excess of the one-month Norwegian interbank offered rate (NIBOR), and W^e is the monthly return on MSCI world stock market index (converted to NOK using the end of month exchange rate) in excess of NIBOR. Columns (4)-(6) include three additional risk factors, SMB for size, HML for value, and MOM for momentum, constructed by Bernt Arne Ødegaard based on OSE-listed firms (available at http://finance.bi.no/∼bernt/). Robust standard errors (White estimator) are in parenthesis and significance levels are indicted by *** 1%, ** 5%, * 10%.

	Low shortfall Portfolio (1)	High shortfall Portfolio (2)	Low-High shortfall Portfolio (3)	Low shortfall Portfolio (4)	High shortfall Portfolio (5)	Low-High shortfall Portfolio (6)
α	0.010**	0.004	0.006	0.012***	0.004	0.008*
	(0.004)	(0.006)	(0.004)	(0.004)	(0.006)	(0.005)
W	0.892***	1.064***	-0.173***	0.862***	1.078***	-0.216***
	(0.127)	(0.127)	(0.059)	(0.135)	(0.137)	(0.060)
W (-1)	0.286***	0.256*	0.030	0.300***	0.254*	0.047
	(0.102)	(0.136)	(0.091)	(0.106)	(0.146)	(0.094)
SMB				-0.262*	-0.015	-0.247***
				(0.147)	(0.181)	(0.085)
HML				0.072	0.053	0.019
				(0.078)	(0.142)	(0.128)
MOM				-0.043	0.036	-0.079
				(0.087)	(0.106)	(0.071)
R^2	0.590	0.546	0.064	0.616	0.547	0.160
Observations (months)	75	75	75	75	75	75

Table 7
Effect on Q of closing the female director shortfall from 2001

Panel A reports estimates of the slope coefficient β in the reduced-form regression for Q of firm i in year t:

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

where *Shortfall women*$_t$ is the shortfall of female directors in year t relative to the quota requirement (see Table 1), and θ_i and τ_t are firm and year fixed effects. Panel B reports estimates of the slope coefficient β in the second-stage (instrumental variable) regression:

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

where $\widehat{Shortfall\ women}$ is the fitted value from the following first-stage regression (reported in Panel C):

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

where D_τ is a year dummy and T_1 and T_2 are the beginning end ending years of the regression period, respectively. The sample comprises 239 OSE-listed ASA firms, 2002-2008 (with Q and board information). Standard errors clustered by firm are reported in parenthesis. Stars indicate significance levels *** 1%, ** 5%, and * 10%.

	Regression period (T_1 to T_2): 2002 to 2007 (1)	Regression period (T_1 to T_2): 2002 to 2008 (2)
A: Reduced-form regression for Q (no instrumentation)		
Shortfall women	0.071	0.304
	(0.345)	(0.374)
F-statistic	10.249	23.670
Firm-years	847	1009
B: 2^{nd} stage regression for Q with $Shortfall\ women_{2001}$ in 1^{st} stage		
$\widehat{Shortfall\ women}$	0.804	0.750
	(0.837)	(0.737)
F-statistic	10.988	18.526
Firm-years	699	815
C: 1^{st} stage instrumentation regressions with $Shortfall\ women_{2001}$		
$D_{2003} Shortfall\ women_{2001}$	-0.101	-0.102
	(0.083)	(0.083)
$D_{2004} Shortfall\ women_{2001}$	-0.176*	-0.178*
	(0.099)	(0.098)
$D_{2005} Shortfall\ women_{2001}$	-0.365***	-0.366***
	(0.103)	(0.102)
$D_{2006} Shortfall\ women_{2001}$	-0.533***	-0.537***
	(0.102)	(0.101)
$D_{2007} Shortfall\ women_{2001}$	-0.630***	-0.632***
	(0.106)	(0.105)
$D_{2008} Shortfall\ women_{2001}$	.	-0.714***
	.	(0.096)
F-statistic	68.507	84.789
Firm-years	726	832

Table 8
Effect on Q of using alternative initial values for Shortfall women and regression periods

Panel A reports estimates of the slope coefficient β in the following reduced-form, fixed effect panel regression for Tobin's Q of firm i in year t:

$$Q_{i,t} = \alpha + \beta Shortfall\ women_{i,t-1} + \theta_i + \tau_t + \epsilon_{i,t},$$

where $Shortfall\ women_t$ is the shortfall of female directors relative to the quota requirement (see Table 1) in year t, and θ_i and τ_t are firm and year fixed effects. Panels B-E report estimates of β in the second-stage IV regression:

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

where $\widehat{Shortfall\ women}$ is the fitted value from the following first-stage regression (not tabulated):

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

where D_τ is a year dummy and T_1 and T_2 are the beginning end ending years of the regression period, respectively. In Panel B, the first-stage instrumentation uses $Shortfall\ women_{2002}$. Panel C and D use $Shortfall\ women_{2001}$ and $Shortfall\ women_{2000}$ in the first stage instrumentation, respectively. The sample comprises 227 OSE-listed ASA firms, 2003-2009 (with Q and board information). Standard errors clustered by firm are reported in parenthesis. Stars indicate significance levels *** 1%, ** 5%, and *10%.

	Regression period (T_1 to T_2): 2003 to 2007 (1)	Regression period (T_1 to T_2): 2003 to 2008 (2)	Regression period (T_1 to T_2): 2003 to 2009 (3)
A: Reduced-form regression for Q (no instrumentation)			
Shortfall women	0.001	0.297	0.323
	(0.387)	(0.421)	(0.451)
F-statistic	4.869	23.258	21.468
Firm-years	717	879	1028
B: 2^{nd} stage regression for Q with $Shortfall\ women_{2002}$ in 1^{st} stage			
$\widehat{Shortfall\ women}$	1.192	1.854**	1.910**
	(0.801)	(0.805)	(0.833)
F-statistic	4.504	16.718	15.851
Firm-years	588	717	820
C: 2^{nd} stage regression for Q with $Shortfall\ women_{2001}$ in 1^{st} stage			
$\widehat{Shortfall\ women}$	1.089	0.922	0.689
	(1.003)	(1.077)	(1.236)
F-statistic	4.762	17.175	16.691
Firm-years	578	694	790
D: 2^{nd} stage regression for Q with $Shortfall\ women_{2000}$ in 1^{st} stage			
$\widehat{Shortfall\ women}$	0.807	0.606	0.246
	(0.877)	(1.034)	(1.199)
F-statistic	4.520	19.388	17.224
Firm-years	522	608	683

Table 9
Firm and director characteristics for sample of ASA and Large AS firms, 1998-2013

The table reports bi-annual firm and board characteristics for the expanded sample of listed ASA firms (Panel A), unlisted ASA firms (Panel B), and Large AS (Panel C) over 1998-2013. Columns (1)-(4) show the median and mean revenue and book value of total assets, reported in million 2013 USD and winsorized at the 1% tails. Columns (5)-(7) list the average fraction of female shareholder-elected directors, Chairmen, and CEOs, respectively. Large AS is the top 1% AS stand-alone firms (excluding subsidiaries) by revenue.

	Revenue		Total asset		Fraction of female			
Year	Mean	Median	Mean	Median	Directors	Chairmen	CEOs	*N*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	
A: Sample of listed ASA firms								
1998	409	96	756	144	0.033	0.016	0.027	196
2000	347	59	727	125	0.046	0.011	0.017	193
2002	451	80	865	121	0.068	0.013	0.033	160
2004	506	74	863	117	0.144	0.019	0.019	155
2006	497	75	1,001	189	0.283	0.017	0.023	175
2008	556	151	1,279	375	0.403	0.031	0.026	193
2010	577	119	1,229	275	0.414	0.058	0.029	174
2012	665	146	1,303	361	0.407	0.101	0.031	159
2013	699	163	1,373	266	0.416	0.088	0.047	150
B: Sample of unlisted ASA firms								
1998	46	3	133	6	0.023	0.004	0.017	247
2000	49	3	107	7	0.028	0.008	0.022	387
2002	91	3	226	7	0.046	0.016	0.044	390
2004	99	5	256	11	0.073	0.030	0.037	334
2006	84	5	338	13	0.203	0.035	0.049	289
2008	117	4	511	21	0.398	0.104	0.070	202
2010	133	8	555	25	0.406	0.103	0.112	155
2012	241	14	778	50	0.382	0.088	0.101	91
2013	229	17	787	87	0.371	0.151	0.096	86
C: Sample of Large AS firms								
1998	174	78	180	48	0.074	0.033	0.023	918
2000	132	56	169	38	0.082	0.025	0.034	943
2002	150	60	197	42	0.093	0.034	0.030	963
2004	173	69	219	46	0.108	0.045	0.043	1,003
2006	216	90	280	67	0.130	0.042	0.050	975
2008	340	143	497	116	0.131	0.041	0.052	1,019
2010	275	114	431	99	0.140	0.057	0.067	1,000
2012	274	116	442	103	0.140	0.059	0.070	1,101
2013	293	118	479	103	0.132	0.055	0.072	1,158

Table 10
Estimates of post-quota changes in operating profitability

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

$$\mathrm{ROA}_{i,t} = \alpha + \gamma_1 \mathrm{ASA\ x\ Post}_{i,t} + \gamma_2 \mathrm{Post}_{i,t} + \gamma_3 \mathrm{ASA}_{i,t} + \gamma_4 \mathbf{x}_{i,t} + \epsilon_{i,t}.$$

The dependent variable ROA is the ratio of operating profit (EBIT) to the book value of assets. The three regressions use different sample periods and definitions of the post-quota period (indicated by the dummy variable *Post*). In column (1), *Post*2008 indicates a post-quota period starting in 2009 (the first full year of quota compliance), and the estimation period is 1998-2013 excluding the quota-implementation years 2002-2008 (see Table 2). Columns (2)) and (3) follow Matsa and Miller (2013) and use a post-quota period that starts in 2007, indicated by *Post*2006. The estimation period in column (2) is 2003-2009, also as in Matsa and Miller (2013), and column (3) adds the post-crisis years 2010-2013 to the estimation period. The vector $\mathbf{x}_{i,t}$ contains the following control variables: total number of shareholder representatives on the board (*Board size*), natural log of firm age (*Firm age*), natural log of book value of total assets (*Total assets*), ratio of book value of total debt to total assets (*Leverage*), percent shares of the *Largest owner*, a dummy indicating that the ASA firm is OSE-listed (*Listed*), and industry sector dummies. The sample comprises 866 ASA and 3,213 Large AS, 1998-2013, excluding financial firms and Large AS that are 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%.

	Sample period 1998-2001 and 2009-2013 (1)	Sample period 2003-2009 (2)	Sample period 2003-2013 (3)
ASA x Post2008	0.030*		
	(0.018)		
Post2008	-0.026***		
	(0.005)		
ASA x Post2006		-0.034**	-0.020
		(0.014)	(0.013)
Post2006		-0.030***	-0.035***
		(0.004)	(0.004)
ASA	-0.198***	-0.139***	-0.133***
	(0.017)	(0.016)	(0.016)
Board size	-0.008***	-0.006***	-0.006***
	(0.002)	(0.002)	(0.002)
Firms age	0.014***	0.010***	0.010***
	(0.003)	(0.003)	(0.002)
Total assets	0.031***	0.028***	0.026***
	(0.003)	(0.003)	(0.003)
Leverage	-0.146***	-0.113***	-0.109***
	(0.021)	(0.024)	(0.020)
Largest owner	0.024**	0.013	0.009
	(0.010)	(0.011)	(0.010)
Listed	0.048**	0.026	0.005
	(0.020)	(0.020)	(0.019)
Constant	-0.171***	-0.145***	-0.132***
	(0.039)	(0.045)	(0.041)
Sector controls	Yes	Yes	Yes
R^2	0.243	0.189	0.176
Observations (firm-years)	9,053	7,797	11,140

Table 11
Reasons for exit from the ASA legal form

Column (1) lists the annual population of non-financial ASA, 2001-2007. Columns (2)-(4) lists the number of firm firms in year t in Column (1) that for various reasons exit the ASA classification in the following year, $t+1$. The possible reasons for the exit are inferred as they are not stated explicitly in publicly available sources. In Column (2), the reason for the exit is merger or forced or voluntary liquidation—all cases where the firm does not reappear as an AS. Column (3) shows the number of conversions from ASA to AS due to domestic or foreign acquisition and where the firm continues as subsidiary AS of the parent. Column (4) lists the number of firms converting from ASA to AS for all other reasons (generally unavailable in public data sources but may hypothetically include circumvention of the gender quota). The sample consists of 277 unique listed ASA (Panel A) and 456 unique unlisted ASA (Panel B). Financial firms are excluded since they were required to be ASA until 2007.

		Inferred reasons for the exit from ASA in year $t+1$		
	Number of ASA at year-end t (1)	Exit from ASA *without* conversion to AS, after forced or voluntary liquidation or merger (2)	Conversion from ASA to AS as a subsidiary after acquisition (3)	Conversion from ASA to AS for all other reasons (unspecified) (4)
A: Listed ASA				
2001	157	6	4	0
2002	148	8	4	0
2003	140	3	8	0
2004	144	7	1	0
2005	161	12	4	0
2006	166	7	12	0
2007	195	11	7	0
B: Unlisted ASA				
2001	288	24	1	27
2002	258	29	6	19
2003	221	10	4	19
2004	210	31	14	19
2005	161	12	3	25
2006	153	15	12	27
2007	126	8	7	12

Table 12
The likelihood of an unlisted ASA converting to AS

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

$$\text{Convert}_{i,t} = \alpha + \gamma_1 \text{Shortfall women}_{i,t-1} + \gamma_2 \mathbf{X}_{i,t} + \epsilon_{i,t}.$$

In column (1) and (2), the dependent variable *Convert* takes the value of one if ASA firm i is an AS in year $t+1$ and zero otherwise. In column (3) and (4), we follow Bøhren and Staubo (2014) and let *Convert* take the value of one for *all* ASA firm-years of a firm that converts at some point during the sample period. *Shortfall women* is the quota-induced fraction of female director shortfall. The vector $\mathbf{X}_{i,t}$ contains the following control variables: total number of shareholder representatives on the board (*Board size*), natural log of firm age (*Firm age*), natural log of book value of total assets (*Total assets*), ratio of operating profit (EBIT) to book value of total assets (*ROA*), ratio of book value of debt to total assets (*Leverage*), and percent shares of the (*Largest owner*). All estimations include industry sector dummy variables, while Column (2) and (4) also include year fixed effects. The sample comprises 480 firm-years for 144 firms that convert (*Convert=1*) and 441 firm-years for 127 that do not convert (*Convert=0*), for a total of 821 firm-years, 2001-2007. These are the firms that convert to AS in 2002-2008 for reasons other than liquidation and acquisition (Column 4 of Table 11). Firms drop out of the sample as they convert. Standard errors clustered by firm are reported in parenthesis. Stars indicate significance levels *** 1%, ** 5%, and * 10%.

	Dependent variable Convert=1 if firm is ASA at year-end t and AS at year-end *t+1*		Dependent variable Convert=1 if ASA firm converts to AS at some point during the sample period (back-filling)	
	(1)	(2)	(3)	(4)
Shortfall women	0.271	0.557	3.724***	1.513
	(0.595)	(0.738)	(0.814)	(1.024)
Board size	0.007	0.021	-0.112	-0.141
	(0.081)	(0.085)	(0.110)	(0.113)
Firms age	0.092	0.094	0.155	0.091
	(0.095)	(0.100)	(0.156)	(0.173)
Total assets	-0.120**	-0.133**	-0.228**	-0.194*
	(0.058)	(0.062)	(0.096)	(0.101)
ROA	-0.408	-0.530*	-0.462	-0.560
	(0.287)	(0.296)	(0.394)	(0.404)
Leverage	0.351	0.427	0.928**	1.003**
	(0.292)	(0.313)	(0.426)	(0.459)
Largest owner	1.401***	1.439***	1.638***	1.476**
	(0.328)	(0.350)	(0.558)	(0.581)
Constant	-2.135***	-2.588***	-0.413	-1.785*
	(0.610)	(0.724)	(0.938)	(1.054)
Sector controls	Yes	Yes	Yes	Yes
Year fixed effects	No	Yes	No	Yes
Pseudo R^2	0.053	0.078	0.195	0.254
Firm-years	821	821	821	821

Table 13
Director network power and gender power gap, 1998-2013

The table reports the annual mean and median director network power for male and female directors serving on at least one ASA board in a given year. In columns (5) and (6), *Power gap* is one minus the ratio of mean and median, respectively, female network power to male network power. Network power is the PageRank centrality score for each year and director in the network, scaled by the maximum score in that year. A higher value indicates greater network power. See the Appendix for details on PageRank. The network power measure is constructed using all directors and CEOs on all ASA and Large AS (defined the top 1% stand-alone AS), 1998-2013. The sample comprises 96,251 directorship-years for 19,206 unique directors in 1,126 unique ASA and 3,473 unique Large AS firms (with board information).

	Male director power			Female director power			Power gap	
Year	Mean (1)	Median (2)	*N*	Mean (3)	Median (4)	*N*	Mean (5)	Median (6)
1998	0.146	0.128	1,534	0.121	0.116	58	0.173	0.089
1999	0.147	0.131	1,693	0.134	0.120	68	0.089	0.084
2000	0.177	0.156	1,929	0.160	0.151	85	0.095	0.031
2001	0.174	0.154	1,937	0.161	0.147	104	0.071	0.047
2002	0.181	0.160	1,773	0.177	0.155	116	0.021	0.030
2003	0.168	0.151	1,624	0.161	0.135	133	0.039	0.103
2004	0.172	0.152	1,612	0.161	0.148	201	0.067	0.029
2005	0.216	0.192	1,427	0.204	0.182	317	0.054	0.049
2006	0.195	0.175	1,340	0.188	0.169	432	0.035	0.035
2007	0.189	0.169	1,145	0.186	0.159	614	0.018	0.059
2008	0.193	0.171	919	0.189	0.168	581	0.025	0.021
2009	0.232	0.202	808	0.223	0.202	506	0.038	0.000
2010	0.237	0.203	775	0.223	0.198	508	0.058	0.023
2011	0.235	0.200	714	0.229	0.193	452	0.025	0.037
2012	0.269	0.226	631	0.257	0.213	391	0.044	0.056
2013	0.270	0.224	593	0.251	0.208	377	0.073	0.073

Table 14
Determinants of board turnover

The table reports the coefficient estimates from the following regression:

$$\text{Board turnover}_{i,t} = \alpha + \gamma_1 \text{ASA x Post2008}_{i,t} + \gamma_2 \text{Post2008}_{i,t} + \gamma_3 \text{ASA}_{i,t} + \gamma_4 \mathbf{x}_{i,t} + \epsilon_{i,t},$$

Board turnover is the fraction of directors of firm i at year-end t that have left at year-end $t+1$. The dependent variable is turnover for the whole board in column (1), for male directors in column (2), and for female directors in column (3). *Post2008* indicates the period 2008-2013. The vector $\mathbf{x}_{i,t}$ contains the following variables: *Implementation* indicates years 2001-2007, average *Female board power*, average *Board power*, dummy variables indicating *Female CEO* and *Female chair*, number of shareholder-elected directors (*Board size*), natural log of *Firm age* and book value of *Total assets*, ratio of EBIT to book value of assets (*ROA*), ratio of book value of debt to assets (*Leverage*), share fraction of the *Largest owner*, and a dummy indicating that the firm is *Listed*. All estimations include industry sector dummies. The sample comprises 855 ASA and 3,060 Large AS, excluding financial firms and Large AS registered as ASA at some point during our sample period, 1998-2013. Standard errors clustered by firm in parenthesis. Stars indicate significance levels *** 1%, ** 5%, and * 10%.

	Total turnover (1)	Male turnover (2)	Female turnover (3)
ASA x Post2008	0.023	0.025	-0.001
	(0.015)	(0.016)	(0.042)
Post2008	-0.040***	-0.044***	-0.023
	(0.008)	(0.008)	(0.019)
ASA x Implementation	0.031**	0.035***	0.016
	(0.012)	(0.012)	(0.041)
Implementation	-0.033***	-0.034***	-0.043**
	(0.007)	(0.007)	(0.018)
ASA	0.037***	0.037***	0.034
	(0.013)	(0.013)	(0.043)
Female board power	0.032	0.022	0.088
	(0.023)	(0.024)	(0.064)
Board power	-0.113***	-0.090**	-0.212**
	(0.042)	(0.044)	(0.101)
Female CEO	-0.017	-0.013	-0.043***
	(0.011)	(0.013)	(0.015)
Female chair	0.001	0.001	-0.012
	(0.012)	(0.014)	(0.016)
Board size	0.026***	0.026***	0.024***
	(0.002)	(0.002)	(0.003)
Firms age	-0.013***	-0.014***	-0.016***
	(0.002)	(0.002)	(0.004)
Total assets	-0.004*	-0.003	-0.010***
	(0.002)	(0.002)	(0.004)
ROA	-0.106***	-0.109***	-0.060*
	(0.015)	(0.015)	(0.034)
Leverage	0.038***	0.031***	0.119***
	(0.012)	(0.012)	(0.024)
Largest owner	0.072***	0.070***	0.062***
	(0.009)	(0.009)	(0.018)
Listed	0.043***	0.039***	0.060***
	(0.011)	(0.011)	(0.022)
Constant	0.095***	0.058**	0.149***
	(0.027)	(0.027)	(0.053)
Sector controls	Yes	Yes	Yes
R^2	0.097	0.091	0.050
Observations (firm-years)	13,621	13,594	5,659

Appendix Table 1
Abnormal returns on February, 22, 2002, following Ahern and Dittmar (2012) (AD)

The table reports abnormal returns (in %) around February 22, 2002, using the event study methodology and sample of 94 OSE-listed ASA firms in AD. AD define abnormal return as the return of the OSE firm minus that of U.S. listed companies matched on industry:

$$AR^{AD}_{i,OSE}(-2,2) = r_{i,OSE}(-2,2) - r_{match,US}(-2,2).$$

Industry classifications are from the Global Industry Classification Standard, while stock returns are from CRSP for U.S. firms and from Compustat Global for Norwegian firms. We use the same data sources here. The number of female directors is per year-end 2001, identified here using the *Brønnøysund* register, yielding 69 firms with no female directors (while AD reports only 68 such firms). In Panel A, the original AD sample includes firms with at least one price observation in their five-day window (-2,2). p-values are in parenthesis, with significance levels *** 1%, ** 5% and * 10%.

	All firms in AD (1)	AD firms with 0 female directors (2)	AD firms with ≥1 female director (3)	Difference (2) - (3)
A: Original AD estimates (their Table III). Event window (-2,2)				
Mean	-2.573***	-3.547***	-0.024	-3.523***
p-value	(0.00)	(0.00)	(0.98)	(0.01)
Number of firms	94	68	26	
B: Our replication of AD with event window (-2,2)				
Mean	-2.817***	-3.770***	-0.188	-3.582**
p-value	0.000	0.000	0.852	0.012
Number of firms	94	69	25	
C: Our replication of AD with event window (-1,0)				
Mean	-0.820	-1.312	0.543	-1.854
p-value	0.201	0.113	0.485	0.102
Number of firms	83	61	22	

Appendix Table 2
Selection of the sample of Large AS, 1998-2013

The table reports the annual number of AS firms that pass cumulative sample filters applied in each year. *Large AS* is the the top 1% percent stand-alone AS firms by revenue. Column (1) reports the population of active AS firms, excluding those listed as inactive. Column (2) further eliminates subsidiary firms that do not report consolidated accounts. Column (3) requires the firms in column (2) to have strictly positive revenue. In column (4), we further eliminate firms that do not pass all of these following criteria: Total assets>0, Long term assets≥ 0, Current assets≥ 0, Long term debt≥ 0, Current debt≥ 0, Current assets$\geq$Cash, and Assets$\geq$Working capital (current assets minus current debt). Finally, Column (5) shows the 1% largest AS of those in column (4), referred to as Large AS. The data is from the Norwegian Corporate Registry. If available, we use consolidated accounts.

Year	All active AS firms (1)	Eliminate subsidiary firms (2)	Require Revenue>0 (3)	Requiring other financial information (4)	Sample of Large AS (5)
1998	125,217	104,835	92,419	91,882	918
1999	130,440	106,322	91,936	91,339	913
2000	135,945	110,296	95,007	94,354	943
2001	138,417	111,031	95,497	94,934	949
2002	139,570	112,430	96,890	96,347	963
2003	135,279	110,825	97,225	96,787	967
2004	142,834	115,335	100,858	100,350	1,003
2005	146,799	104,480	89,163	88,792	887
2006	172,680	131,835	98,108	97,654	975
2007	185,332	133,266	98,574	98,128	981
2008	191,236	139,933	102,340	101,916	1,019
2009	192,601	135,815	99,356	98,920	989
2010	196,057	137,102	100,510	100,087	1,000
2011	201,559	140,451	102,629	101,963	1,019
2012	215,416	151,799	111,152	110,167	1,101
2013	227,841	160,307	117,103	115,821	1,158
Mean	167,326	125,379	99,298	98,715	987

Appendix Table 3
Industry distribution of the sample firms

The table reports the fraction of firm-years in each industry sector, 1998-2013, for the population of 409 listed and 888 unlisted ASA firms and the sample of 3,506 Large AS firms (with sector information).

Industry sector	Listed ASA	Unlisted ASA	Large AS	All firms
Agriculture	0.027	0.024	0.026	0.025
Offshore/Shipping	0.246	0.067	0.099	0.111
Transport	0.016	0.012	0.039	0.031
Manufacturing	0.159	0.071	0.143	0.132
Telecom/IT/Tech	0.180	0.118	0.030	0.065
Electricity	0.013	0.009	0.044	0.034
Construction	0.062	0.051	0.186	0.146
Wholesale/Retail	0.060	0.074	0.227	0.179
Finance	0.063	0.383	0.059	0.118
Other services	0.174	0.191	0.147	0.158

A Definition of PageRank network centrality

Suppose a director network consists of N nodes, where each node is a director on a board. In this network, all connections are reciprocal: if director i connects to director j, then j also connects to i. Define the network's *adjacency matrix* as the $N \times N$ matrix $\mathbf{A}$ where entry a_{ij} is equal to 1 if node i is connected to node j, else zero. Because connections are reciprocal, $\mathbf{A}$ is symmetric. Define *Degree centrality* as the number of other nodes directly connected to a node. The vector containing degree centrality for each node i in the network is the vector of row sums of the adjacency matrix:

$$degree(i) = \sum_{j \neq i} a_{ij}$$

Moreover, let λ_{dom} denote the dominant (largest) eigenvalue of $\mathbf{A}$. Define the vector containing the *eigenvector centrality* of each node in the network as the vector $\mathbf{x}$ such that

$$\mathbf{x} = \lambda_{dom}^{-1} \mathbf{A} \mathbf{x}.$$

Intuitively, in addition to counting first order connection like degree centrality, eigenvector centrality also takes into account the relative importance of each of these connections.

PageRank, the algorithm underlying the Google search engine (Page, Brin, Motwani, and Winograd, 1999), is a modified eigenvector centrality measure for disconnected networks. Two networks are disconnected if there are no connections between them—they are completely isolated from each other. In this case, eigenvector centrality produces a non-zero centrality score for one network with zero weight on the other. Intuitively, PageRank is a modified eigenvector centrality measure which allows for some degree of connections between otherwise disconnected networks. The analogy for disconnected director networks would be to allow for non-board connections between directors.

The first step in the computation of PageRank is to normalize each column in the adjacency matrix by the column sum, so that each column in the normalized adjacency matrix $\mathbf{A}^*$ sums to one. Second, a non-zero random transition matrix $\mathbf{B}$ is constructed, where each entry $b_{ij} = 1/N$. Third, these two matrices are combined linearly to produce a new matrix

$$\mathbf{C} = [1 - \delta]\mathbf{A}^* + \delta \mathbf{B},$$

where δ is the weight given to the random transition matrix (typically, and in this paper, $\delta = 0.15$).

Next, we find the eigenvector of the dominant eigenvalue of this new combined matrix $\mathbf{C}$ as the vector $\mathbf{x}$ such that

$$\mathbf{x} = \lambda_{dom}^{-1} \mathbf{C} \mathbf{x} = \lambda_{dom}^{-1} \Big([1 - \delta]\mathbf{A}^* + \delta \mathbf{B} \Big) \mathbf{x}$$

It can be shown that largest eigenvalue of the combined matrix $\mathbf{C}$ is 1 (Perron-Frobenius theorem), and therefore we are looking for the $\mathbf{x}$ such that

$$\mathbf{x} = \Big([1 - \delta]\mathbf{A}^* + \delta \mathbf{B} \Big) \mathbf{x}.$$

$\mathbf{x}$ can be interpreted as the solution to a system of linear equations. Finally, we normalize this vector by the sum of the elements in the vector to produce the PageRank of each node in the overall network.