Board Gender-Balancing and Firm Value

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Abstract

In 2003, Norway was the first country to mandate gender-balanced corporate boards. Widely cited research concludes that the gender quota caused large declines in market values of companies listed on the Oslo Stock Exchange. We reverse this conclusion after replicating, correcting and extending the earlier analysis. The correct conclusion is that quota-induced changes in market valuations were economically and statistically insignificant. We corroborate this conclusion with evidence that boards managed to preserve directors' large-firm CEO experience without increasing board seat concentration or director busyness. Furthermore, we show that operating profitability did not decline after quota compliance, and that corporate conversions to an unregulated legal form were independent of the quota-induced female director shortfall. Overall, the best available econometric methodology produces evidence strongly suggesting a value-neutral effect of the Norwegian gender quota - as expected when the pool of qualified female directors is large.

Keywords: Gender quota, director independence, valuation effect, long-run performance, corporate conversion, busy directors, director network power

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BOARD GENDER-BALANCING AND FIRM VALUE*

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July 13, 2018

Abstract

In 2003, Norway was the first country to mandate gender-balanced corporate boards. Widely cited research concludes that the gender quota caused large declines in market values of companies listed on the Oslo Stock Exchange. We reverse this conclusion after replicating, correcting and extending the earlier analysis. The correct conclusion is that quota-induced changes in market valuations were economically and statistically insignificant. We corroborate this conclusion with evidence that boards managed to preserve directors' large-firm CEO experience without increasing board seat concentration or director busyness. Furthermore, we show that operating profitability did not decline after quota compliance, and that corporate conversions to an unregulated legal form were independent of the quota-induced female director shortfall. Overall, the best available econometric methodology produces evidence strongly suggesting a value-neutral effect of the Norwegian gender quota—as expected when the pool of qualified female directors is large.

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"[Norway's gender] quota caused a significant drop in the stock price...and a large decline in Tobins Q over the following years...The quota led to younger and less experienced boards,.., and deterioration in operating performance."
— Kenneth R. Ahern and Amy Dittmar, *Quarterly Journal of Economics*, 2012 (abstract).

1 Introduction

In late 2003, Norway passed a law requiring public limited companies (ASA) to have gender-balanced boards. Two years later, firms were given two years to comply—with mandatory compliance by year-end 2007—or face forced liquidation. The quota law was driven by gender politics unrelated to firm performance (*Odelstingsproposisjon* 97, 2002-2003). For this reason, passage of the quota provides a rare, quasi-experimental setting for assessing causal effects of exogenous changes in board composition on firm value (Adams, Hermalin, and Weisbach, 2010). Several other European countries (Belgium, France, Germany, Iceland, Italy, the Netherlands, Portugal, and Spain) have since followed suit and implemented gender quota laws with varying fractions of female directors and different penalties for non-compliance. Thus, understanding the economic effects of Norway's pioneering quota is of interest well beyond Norway.

Ahern and Dittmar (2012) document that the post-reform female board members were professionals with better education than male directors. Also, Bøhren and Staubo (2015) find evidence of increased director independence, which is commonly viewed as important for shareholder value maximization (Jensen, 1993). Moreover, Bertrand, Black, Jensen, and Lleras-Muney (2017) show that the new female ASA board members were more qualified than their female predecessors. Add to this potential benefits of increased board diversity (Kim and Starks, 2016; Bernile, Bhagwat, and Yonker, 2018), and it is possible that the quota law could lead to higher firm value. This value-enhancing hypothesis raises, of course, the question of why large shareholders did not already elect more diverse boards. While this paper is not about gender politics, one possibility is that forced gender-balancing may have been required to unlock such benefits, even in the relatively egalitarian Norwegian society.

On the other hand, based on biographical information in annual reports, Ahern and Dittmar (2012) find that publicly traded ASA experienced a decline in board CEO experience. There is evidence elsewhere in the governance literature that director CEO experience is valued by investors (Fich, 2005; Fahlenbrach, Low, and Stulz, 2010). One reason is that such experience enhances the quality of the board's advisory function. Since the new female directors are less experienced, they may be overly focused on monitoring

(Adams and Ferreira, 2009; Bøhren and Staubo, 2015). Overall, it is an empirical issue whether negative effects of a reduction in board CEO experience and excessive monitoring are sufficient to outweigh potential beneficial valuation effects of greater director education and independence.[1]

In the most influential prior study of the Norwegian quota, Ahern and Dittmar (2012) report large negative overall valuation effects of the quota (above quote). Most striking is their estimate of "value losses upwards of 20% for the firms with large constraints" (i.e, all-male boards) that they attribute to an influx of younger and less experienced female directors.[2] To put this in perspective, it implies that the quota may have caused losses on par with dramatic corporate events such as major product recalls and severe financial distress (Jarrell and Peltzman, 1985; Hotchkiss, John, Mooradian, and Thorburn, 2008). It also suggests that forced gender-balancing may have been a much more costly board restriction than the 2002 US Sarbanes Oxley Act (SOX), which mandates expensive internal control systems (Chhaochharia and Grinstein, 2007; Duchin, Matsusaka, and Ozbas, 2010). This notwithstanding the fact that, in contrast to SOX, the Norwegian quota law narrowly regulates gender balancing only. It caused the typical ASA board in 2001 to replace two of its five male directors with females over a six-year period.

In this paper, we resolve the puzzling disconnect between the limited extent of the actual ASA board changes and Ahern and Dittmar (2012)'s large negative valuation estimates. Our analysis includes replicating, correcting, and expanding the empirical analysis of Ahern and Dittmar (2012) and of subsequent influential studies of the gender quota. We draw five main conclusions. Fist and foremost, there is no evidence of a statistically significant impact of the quota law on the equity values of ASA listed on the Oslo Stock Exchange (OSE). Since we include a replication of Ahern and Dittmar (2012)'s own estimates—using their methodology and sample of firms—this conclusion is not driven by differences in either data or programming. Rather, our estimates and overall conclusion of neutral valuation effects replace theirs.

Second, we show that the evidence in Matsa and Miller (2013) of negative operating profitability following quota compliance is short-term. There is no change in operating profitability following quota compliance once we extend the sample period beyond the recent financial crisis. Third, notwithstanding

[1]There is also a growing empirical literature arguing that female executives are more risk averse and less confident than males (Huang and Kisgen, 2013; Levi, Li, and Zhang, 2014). If this is generally true, forced gender-balancing may lower board risk-taking and, therefore, equity value. However, Sila, Gonzales, and Hagendorff (2016) find no relation between board gender diversity and equity risk.

[2]This estimate received intense media coverage at the time: *The Financial Times*: "[the quota caused] a large decline in Tobins Q...over the following years." (August 20, 2011). *The Economist*: "[Norway's gender quota] led to large numbers of inexperienced women being appointed to boards, and...has seriously damaged those firms' performance." (July 21, 2011). *The Wall Street Journal*: "[the quota law] damaged shareholder value in the companies affected." (June 11, 2012).

Ahern and Dittmar (2012)'s evidence of a decline in board CEO experience, we show that board *large-firm* CEO experience—which excludes CEO experience in tiny firms (zero to two employees)—did *not* decline. Fourth, we document that quota compliance was obtained without increasing director busyness, suggesting that there was no shortage of qualified female directors. Fifth, as Nygaard (2011) and Bøhren and Staubo (2014), we study legal conversions from ASA to the unregulated private limited company (AS) form that occurred before, during, and after the quota compliance period. Firms do not state the reason for converting. However, unlike the earlier research, we show that conversions are statistically unrelated to the quota-induced female director shortfall in the cross-section.

We begin our empirical analysis with a greatly expanded estimation of the valuation impact of major quota-related news events on the portfolio of OSE-listed ASA. As the news events affect all firms at the same points in calendar time, we form a portfolio of listed ASA to efficiently incorporate cross-dependence in stock returns (Schwert, 1981; Kolari and Pynnönen, 2010). The conclusion from this portfolio estimation is that the market reaction to quota news events is insignificant on average. A similar conclusion holds for the single quota news event used by Ahern and Dittmar (2012), whether based on their estimation procedure and sample of firms, or our larger data set.[3] Moreover, cross-sectional regressions with individual ASA announcement returns as dependent variable show a statistically insignificant impact of the quota-induced female director shortfall.

We also follow Ahern and Dittmar (2012) and, using their two-stage instrumental variable (IV) approach, estimate effects of the quota on Tobin's Q over the legislative period. We observe that almost half of the listed ASA began voluntarily increasing the fraction of female directors following the initial public quota announcement in February of 2002. The first IV stage generates an instrument designed to eliminate effects of such firm-specific (endogenous) compliance timing from 2002 to 2008. The instrument replaces the actual evolution of the firm's female director shortfall with the market-wide compliance trend from an exogenous, firm-specific starting point in 2001. In the second stage, Tobin's Q is regressed on the instrumented fraction of female director shortfall, controlling for firm and year fixed effects.

Importantly, we conclude from the IV tests that the female director shortfall does *not* impact Tobin's Q, whether using the earlier sample or our own. Ahern and Dittmar (2012)'s conclusion of a value loss of "upwards of 20%" in Tobin's Q holds only when using year-end 2002 as basis for the first-stage instrumentation. However, since many firms began complying with the expected quota already at the 2002

[3] We thank Kenneth Ahern for providing a list of their sample firms.

annual shareholder meeting, it is necessary to push back the starting point to 2001 to avoid endogeneity bias in the second stage. We demonstrate that the earlier evidence of a significant impact of the quota on Q becomes statistically insignificant when we eliminate this bias. In addition, when we replace the fraction of female director with the *number* of missing females, which is unaffected by simultaneous changes in board size, the quota has no significant impact on Q, whether using 2001 or 2002 as starting point in the instrumentation. Thus, the evidence strongly suggests that the quota was value-neutral.

Second, we follow Matsa and Miller (2013) and use a difference-in-difference test for quota effects on operating profitability, measured by return on assets (ROA). This approach exploits the information in all ASA (not just the OSE-listed ones), using the performance of the 1% largest (by revenue) unregulated AS as control sample. Matsa and Miller (2013) conclude that ROA drops for ASA after mandated quota compliance, ending their sample in 2009. We reach a similar conclusion when ending the panel data estimation in 2009 for a substantially larger sample of ASA (741 treated firms versus 104 in their study). However, we find no significant decline in post-quota operating profitability of ASA when we extend the sample period to 2013. Thus, the decline in ROA reported by Matsa and Miller (2013) appears to be a phenomenon limited to the financial crisis years.

Third, we document various board changes as an indirect validation of the insignificance of our valuation and profitability estimates. Our most important finding here concerns the change in board CEO experience. Using biographical information from company annual reports on directors' CEO experience, which includes positions in directors' wholly-owned firms, Ahern and Dittmar (2012) report in their online appendix that, in 2001, 67% of male and as much as 53% (!) of female directors in listed ASA had CEO experience. In 2009, the percentage is higher (70%) for males and lower (39%) for females, with a drop in overall board CEO experience. This is the basis for their claim that the quota deprived boards of valuable experience.

However, we reach a substantially different conclusion than Ahern and Dittmar (2012) also on this point. To place their high percentage of female director CEO experience in context, note first that only 5% of all ASA (4% of listed ASA) had a female CEO in 2001. We confirm below that their measure of CEO experience is dominated by tiny AS with zero to two employees that typically have few operational activities of relevance for listed ASA. When we restrict our measure of CEO experience to include ASA and the 1% largest AS only, our difference-in-difference tests show no significant decline in board CEO experience following quota compliance, further validating our conclusion that the quota had a value-neutral effect on OSE-listed firms.

Additional indirect validation comes from our finding of a decline in ASA director busyness over the post-quota period, and that overall (male and female) board seat concentration did not increase. When coupled with our evidence of a neutral valuation effect of the quota, this additional information, which is also new to the literature, is consistent with a sufficiently large supply of qualified female directors. Also, shareholders on average found it unnecessary to expand board size to make room for new female directors without having to dismiss males. Finally, we show that legal conversions from ASA to the unregulated AS corporate form are unrelated to the firm's quota-induced female director shortfall. Once we add year fixed effects to correct for the time trend in the fraction of female directors that influences the main model specification of Bøhren and Staubo (2014), this conclusion holds also for the conversion likelihood that they estimate.

The rest of the paper is organized as follows. Section 2 reports on our quota event study, followed by Section 3 on changes in Tobin's Q and operating profitability. We present our evidence on changes in board CEO experience in Section 5, and on other board changes in Section 6. Section 7 concludes the paper.

2 Quota-induced valuation effects

We begin our analysis of the valuation impact of gender-balancing by estimating the stock market reaction to major news events that increased the likelihood of a quota law. We first explain how the quota varies with board size and define the quota's legislative time line, before presenting our portfolio estimates of abnormal stock returns to major quota news announcements. The section ends with a detailed explanation of why our event study results and overall conclusion differ substantially from the prior evidence.

2.1 The gender quota and board size

In Norway, director elections and decisions on board size are substantially influenced by large shareholders. In listed ASA, the largest shareholder owns on average 37% of the outstanding shares (median 31%). Director nominations occur via an independent nomination committee, which stays in close contact with the largest blockholders. It is also worth mentioning that the quota is a gender quota—not a female quota: It holds for both male and female directors. Also, employees appoint their own directors (up to one-third of the total board) and, since these directors follow other rules and are not regulated by the quota, we follow prior studies and exclude them from our analysis.

Table 1 shows that the mandated fraction of female directors depends on board size (the number of shareholder-elected directors) and ranges from 33% to 50%.For example, an ASA board with five directors must have a minimum of 40% female directors (two women), while the female requirement is 50% for boards with four and six members (two and three women, respectively). Columns (3)-(5) of Table 1 list the fraction of new female directors required by the quota for a given board size, labelled $Shortfall$. Throughout the paper, we use $Shortfall$ to measure the quota constraint facing individual ASA. Since the required fraction of female directors vary with board size, the firm can affect $Shortfall$ either by replacing male directors or by changing board size. As shown, increasing the board from four to five members allows a firm to appoint two females while retaining three (rather then two) males, while reducing board size from six to five directors allows the firm to appoint two (rather than three) females while retaining three males.

In 2001, no ASA board was gender-balanced, and as shown by Figure 1, the average ASA had a five-member all-male board (data source: the national *Brønnøysund Register Centre* and (Berner, Mjøs, and Olving, 2013)). Hypothetically, by expanding board size to eight, this ASA could have chosen to retain all five males and appoint three new females to fill the quota. Not a single ASA made this decision, however, and the average board size remained at five directors throughout the sample period.

This evidence is interesting as it suggests that the perceived cost of increasing board size from five to eight directors exceeded the expected benefit of retaining the five male directors. In other words, the cost of expanding board size to eight directors places an upper bound on the expected shareholder-borne cost of the quota. The fundamental empirical question motivating this paper is whether such an upper bound can possibly exceed Ahern and Dittmar (2012)'s estimated loss of upwards of 20% in Tobin's Q, which is unprecedented in literature on board size (Bhagat and Jefferis, 2005; Boone, Field, Karpoff, and Raheja, 2007; Adams, Hermalin, and Weisbach, 2010).

2.2 The quota time line

The gender quota was passed by the Norwegian parliament at the end of 2003. However, at the time, the law was not mandated by Cabinet (which it must be to take effect), and it had a sunset provision after two years: It would expire if ASA voluntarily complied within the two-year period. The appointment of female directors was slow, and so Cabinet mandated the quota in December of 2005. At that point, existing ASA were given two years to comply, while new ASA were required to comply immediately. All ASA were in full compliance by the annual meetings in the spring of 2008.

We focus our event study on the five news event dates listed in Table 2, all of which gradually increased the probability of a quota law. The first event occurred on Friday, February 22, 2002, when the Minister of Industry and Trade took the market by surprise by declaring his support for a mandatory gender quota in a newspaper interview. The declaration contradicted the official policy of his political party in the coalition government. Likely as a result, he publicly retracted his support the very next day (not previously recognized by the literature).

Given the coalition government's opposition to a gender quota, it came as a further surprise when, on International Women's day (March 8, 2002), the government proposed a gender quota legislation (event 2). Since the major opposition party was in favor of a quota, this indicated that the quota had broad political support. The proposal was submitted to Parliament on June 16, 2003 (event 3). It contained a sunset provision, cancelling the quota if firms complied voluntarily within two years. As expected, it was passed by Parliament's lower chamber ("Odelstinget") and upper chamber ("Lagtinget") in late 2003, and was formally amended to Corporate Law on December 19, 2003.

Although many firms immediately began to increase female director representation, the degree of voluntary compliance was ultimately deemed insufficient by the government. Thus, on December 1, 2005, the Prime Minister announced that the quota would be mandated (event 4). This announcement did not, however, fully resolve the question of sanctions. The issue of sanctions was resolved when Cabinet mandated the law on December 9 (event 5), without any special sanction for quota non-compliance. This meant that the sanction would be forced liquidation, which is the ultimate penalty for violating Norwegian corporate law.

2.3 Portfolio estimation of abnormal returns

We explore the five events in Table 2 both individually and jointly using daily stock returns to OSE-listed ASA, 1998-2013. Daily stock returns and number of shares are from the OSE data service *Oslo Børsinformasjon*, accessed through *Børsprosjektet* at the Norwegian School of Economics. Daily stock returns are computed using differences in (log of) daily closing prices, adjusted for splits and dividends. If a closing price is missing, it is replaced by the bid-ask midpoint, if available (this occurs in twenty percent of the trading days).

To maximize test power (Brown and Warner, 1980, 1985), we use the standard two-day event window, which ends with the public announcement date (day 0). Also, since the news events affect all sample firms simultaneously in calendar time, we account for the resulting contemporaneous cross-correlation of stock returns by forming equal-weighted calendar-time portfolios of the OSE-listed ASA. Inferences do not change if we instead use a value-weighted portfolio or a portfolio with weighted-least-squares weights, i.e., weights inversely proportional to the standard deviation of each individual sample firm's residual ε in the return generating process (Eq. (1) below).[4]

For each of the five events, $k = 1, ..., 5$, we estimate the portfolio's daily abnormal return parameter AR_k, using the following return-generating process:

$$r_t = \alpha + \beta r_{wt} + AR_k d_{kt} + \varepsilon_t. \tag{1}$$

Here, r_t is the daily stock return to OSE-listed ASA, converted to USD with the daily exchange rate, in excess of the daily 3-month US Treasury bill, r_{wt} is the daily excess return on the MSCI stock market world index, and d_{kt} is a dummy variable that takes a value of one for each day in the two-day event window (-1,0) and zero otherwise. Thus, the event parameter AR_k is the average daily abnormal portfolio return over the two event days, and the two-day cumulative abnormal return is $CAR_k(-1, 0) = 2AR_k$.[5]

Since the ratio of the cumulative abnormal return and its standard deviation is the same as the ratio of the average abnormal return and its standard deviation, the t-statistic of $CAR_k(-1, 0)$ is $t = \frac{2AR_k}{\sigma_{2AR_k}} = \frac{AR_k}{\sigma_{AR_k}}$, where the estimate of σ_{AR} is conveniently provided by the regression package used to estimate Eq. (1). The regression starts 252 trading days prior to and ends on event date k (day 0), excluding days of prior events, if any. To be included in the portfolio, a firm must have a minimum of 100 return observations and, importantly, an actual return observation on each day in the event window. While not

[4] An alternative to our portfolio formation is to estimate abnormal returns using a system of seemingly unrelated regressions (SUR), with a single OLS regression for each sample firm, and compute standard errors accounting for the residual cross-correlation. We prefer the portfolio approach since the SUR approach simply reproduces our portfolio estimates, while substantially increasing the number of required parameter estimates. Inferences based on the SUR approach are asymptotically identical to those of our portfolio estimation. However, SUR may be less efficient because it increases parameter estimation error (Sefcik and Thompson, 1986). Kolari and Pynnönen (2010) simulate the relative power of the SUR approach for a range of event date clustering.

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

tabulated, our main empirical inferences below are unaffected if we use a simple mean-adjusted return model ($r_t = \alpha + AR_k d_{kt} + \varepsilon_t$), a lead-lag market adjustment to account for non-synchronous trading as in Scholes and Williams (1977) ($r_t = \alpha + \beta_1 r_{w,t-1} + \beta_2 r_{wt} + \beta_3 r_{w,t+1} + AR_k d_{kt} + \varepsilon_t$), or a four-factor model (Fama-French factors and momentum, from Ken French's web site).

Table 3 reports the portfolio estimates of the two-day abnormal return, $CAR_k(-1, 0) = 2AR_k$, along with p-values in square brackets. The abnormal return in the last row is the cumulative abnormal return across all five events, estimated beginning 252 days prior to the first event (February 22, 2002) and ending on the day of the fifth event (December 9, 2005). This estimation, which is added as a check on the sum of the five events, redefines the dummy variable d_t to take a value of one for each of the five two-day event windows. The resulting cumulative abnormal return is therefore $CAR_{1-5}(-1, 0) = 10AR_{1-5}$.

While Column (1) of Table 3 uses a portfolio containing the total sample of listed ASA, columns (2)-(5) split the sample cross-sectionally into various sub-portfolios. Columns (2) and (3) use portfolios of ASA with *High shortfall* and *Low shortfall*, respectively, measured at the year-end preceding each event date. *High Shortfall* are ASA with $Shortfall$ at or above the sample median where, again, $Shortfall$ is the difference between the fraction of female directors required by the quota (Table 1) and the actual fraction of female directors on the board.

With the exception of event 2 (March 8, 2002), none of the five news events in Table 3 generate statistically significant abnormal stock returns. Moreover, none of the five events generate significant abnormal returns for the long-short portfolio in Column (4), which is long in *High shortfall* and short in *Low Shortfall* ASA. This conclusion also holds for *All five events* using $CAR_{1-5}(-1, 0)$, which tests for the joint significance of the five abnormal returns. Moreover, while not tabulated, results with an alternative 3-day event window (-1,1) yield identical statistical inferences, irrespective of the risk adjustment.

The significantly positive abnormal portfolio return in columns (1) and (2) on March 8, 2002 warrants investigation. News searches reveal that, on this day, Parliament approved a plan for the development and operation of the Snøhvit natural gas field in the Barents Sea. Parliament also approved the installation and operation of an onshore plant to process liquefied natural gas from the field. The Snøhvit project, with estimated investments exceeding $6.3 billion, was the first natural gas development in the Barents Sea and Europe's first gas liquefaction project.

Table 3 further shows that the positive abnormal return on March 8, 2002 in columns (1) and (2) is most likely driven by the Snøhvit announcement rather than the gender quota. To see this, note that

Column (5) reports a statistically weak but positive abnormal return to the sub-portfolio of 31 OSE-listed ASA operating in the oil/offshore sector. Moreover, Column (6) shows a similar market reaction to ten OSE-listed foreign-domiciled oil/offshore companies, which are not subject to the quota law. Finally, Column (7) reports a small and statistically insignificant abnormal return to a portfolio long in domestic and short in foreign oil/offshore firms. Thus, the positive abnormal returns appear to be industry-driven, rather than a phenomenon tied to Norwegian incorporation and hence the quota law.

Finally, we perform cross-sectional (OLS) regressions at the firm level in order to test whether the market reaction to quota news events depends on the shortfall of female directors. If the quota constraint is costly, firm i's abnormal return in response to event k, $CAR_{i,k}(-1,0)$, should be more negative the more binding the quota constraint, i.e. negatively correlated with *Shortfall*. For each event k, the regression specification is:

$$CAR_{ik}(-1,0) = \alpha_k + \gamma_{1k} Shortfall_{ik} + \gamma_{2k} \mathbf{X_{ik}} + \kappa_i + u_{ik}. \quad i = 1, \dots, N. \tag{2}$$

In addition to $Shortfall$, the vector of controls $\mathbf{X}$ includes *Largest owner* (percent ownership of the largest shareholder), a dummy indicating *Government control*, *Codetermination* (a dummy indicating that quota-induced females and employee directors together have a majority of the board seats), *Risk* (the firm's daily stock return volatility in the year prior to the event), and *Total assets* (log of book value of total assets). All variables are defined in Table 4. *Codetermination* and *Risk* are meant to capture, respectively, the possibility that new and relatively inexperienced female directors form coalitions with labor representatives on the board and that they may be excessively risk averse, to the detriment of shareholders. The regressions also include industry fixed effects (κ_i) allocating each OSE-listed ASA to one of ten industry sectors. The largest of these sectors is oil/offshore with 25% of the firms, followed by the telecom/technology sector with 18%. The remaining eight sectors are manufacturing (16%), construction (6%), wholesale/retail (6%), finance (5%), agriculture (3%), transportation (2%), electricity (1%), and other services (18%).

The results are shown in Table 5. Panel A uses the fraction of female director shortfall ($Shortfall$), while Panel B uses the number of missing female directors ($Shortfall_{number}$). Importantly, the regressions fail to identify significant effects of either $Shortfall$ or $Shortfall_{number}$ on the event returns for all five events. In sum, the estimates in tables 3 and 5 fail to reject the null hypothesis of a a zero valuation effect of the gender quota for OSE-listed ASA.

While not tabulated, using monthly returns from February 2002 (the first major quota event) through April 2008 (when all firms were in full compliance), we also estimated long-run abnormal return parameter α in the following return generating process:

$$r_t = \alpha + \beta_1 r_{wt} + \beta_2 HML_t + \beta_3 SMB_t + \varepsilon_t, \tag{3}$$

where r_t is the monthly industry-adjusted USD-denominated portfolio stock return, r_{wt} is the monthly return on the MSCI world stock-market index in excess of the 3-month US treasury bill, and HML and SMB are monthly returns to the global value and size factors from Ken French's web site. We focused this analysis on two equal-weighted portfolios: listed ASA with $Zero_{2001}$ (all-male board in 2001) and Pos_{2001} (mixed-gender board in 2001). In this estimation, firms never switch portfolios. The abnormal performance parameter α is statistically insignificant for both portfolios, and also for a third difference-portfolio that is long in $Zero_{2001}$ firms and short in Pos_{2001} firms.

2.4 Ahern and Dittmar (2012)'s event study

Ahern and Dittmar (2012) report significantly negative abnormal stock returns over a five-day window centered on the February 22, 2002, newspaper interview with the Minister of Industry and Trade (our event 1). Since they use a smaller sample of 94 OSE-listed ASA, and because their data sources and estimation methodology differ from ours, we replicate their analysis here. To simplify a comparison of the two studies, Panel A of Table 6 copies the average abnormal return estimates directly from Panel A of Table III in their study. These estimates are averages of firm i's five-day event return, $CAR_i^{AD}(-2,2)$, defined as follows:

$$CAR_i^{AD}(-2,2) = \sum_{\tau=-2}^{2} (r_i - r_{imatch})_\tau \,, \; i = 1, ..., 94. \tag{4}$$

r_i is the return to OSE-listed ASA i on event day τ, and r_{imatch} is the average return to US-listed companies in firm i's Global Industry Classification Standard (GICS) industry. Their return data are from Compustat Global for Norwegian firms and Center for Research in Securities Prices (CRSP) for the matching firms. They report p-values only, and the standard errors for the average $CAR^{AD}(-2,2)$ shown in parentheses are therefore computed by us from their reported p-values. As listed in Column

(1), the sample average $CAR_i^{AD}(-2,2)$ is -2.57% with a reported $p = 0.001$ (implying a standard error of 0.757). Thus, unlike us, they conclude with a large negative market reaction to the February 22, 2002, quota event.

In Panel B of Table 6, we closely replicate the abnormal returns in Panel A by using the same 94 ASA, the $CAR_i^{AD}(-2,2)$ defined in Eq. (4), and return data from Compustat Global and CRSP, but with board data from the *Brønnøysund Register Centre*. We classify 69 of the 94 ASA as having all-male boards in 2001 vs. 68 in Ahern and Dittmar (2012), who use biographical information from annual reports. Our replication yields an average abnormal return estimate of -2.73%, which is close to the -2.57% in Panel A.[6] However, to generate a p-value similar to the one in Panel A, we must assume in Panel B that the individual abnormal return estimates are cross-sectionally independent. To reiterate, although Ahern and Dittmar (2012) do not explicitly make this assumption of cross-sectional independence, it is required in order to replicate their significance levels in Panel A.

In Panel C of Table 6, we generate correct standard errors by re-estimating Ahern and Dittmar (2012)'s five-day industry-adjusted abnormal return using the portfolio time-series approach in Table 3 above. This portfolio estimation starts 252 days prior to February 22, 2002, has the equal-weighted industry-adjusted portfolio return, $r_{p,t}^{-I} \equiv \frac{1}{N}\sum_{i=1}^{N}(r_i - r_{imatch,})_t$, as dependent variable, and uses an event dummy d_t that takes a value of one over the five-day window (-2,2) and zero otherwise. The portfolio estimate is $CAR(-2,2) = 5AR = -2.12\%$, which resembles the -2.57% in Panel A. However, the portfolio estimate of σ_{AR} is 0.650. As a result, the t-statistic, $t = 5AR/\sigma_{5AR} = AR/\sigma_{AR}$, now has an insignificant p-value of 0.516. Following Ahern and Dittmar (2012), we also split the sample by the board gender composition at year-end 2001 into firms with all-male boards ($Zero_{2001}$) in Column (2) and firms with at least one female director (Pos_{2001}) in Column (3). Again, our empirical estimates are similar to theirs, but insignificant.

How large is Ahern and Dittmar (2012)'s understatement of the standard error of their abnormal return estimates? To illustrate the bias, recall that, in panels A and B of Table 6, the standard error of the average five-day abnormal return is computed as $\sigma_{CAR} = \frac{\sigma}{\sqrt{N}}$, where σ is the cross-sectional standard error of the $N = 94$ observations on $CAR_i^{AD}(-2,2)$ in Eq. (4), and where σ falsely assumes cross-sectional independence of the abnormal returns. Assume for illustrative purposes that the N individual

[6] While we use the matching procedure exactly as stated in footnote 9 of Ahern and Dittmar (2012), the discrepancy may nevertheless reflect small differences in the selection of matching firms.

abnormal return variances σ and $N(N-1)$ pairwise return covariances ρ (between firms i and j) are cross-sectionally constant. As in Kothari and Warner (2007), we can then write

$$\sigma_{CAR} = \sqrt{\frac{1}{N}\sigma^2 + \frac{N-1}{N}\sigma^2\rho}, \tag{5}$$

and so the test bias is

$$\frac{\sigma_{CAR}}{\sigma_{CAR}\ independence} = \frac{\sigma_{CAR}}{\sqrt{\sigma^2/N}} = \sqrt{1+(N-1)\rho}. \tag{6}$$

Using $5\sigma_{AR} = 3.250$ from Panel C and $\sigma_{CAR}\ independence = \sigma/\sqrt{94} = 0.780$ from Panel B implies an averageρ of 0.176. While the average pairwise cross-correlation is likely to be higher for the resource-based OSE firms than for firms on the more diversified US stock markets, this estimate is not far from the average pairwise correlation of stock returns of 0.11 reported by de Bodt, Eckbo, and Roll (2018) for the universe of public US manufacturing firms. In sum, their analysis would also have failed to reject the hypothesis of zero abnormal returns had they correctly used standard errors adjusted for cross-dependence in event returns.

2.5 A closer look at AD's event window

Recall from Table 2 that, while the Minister of Industry and Trade made a surprising statement in support of a gender quota on Friday, February 22, 2002, he publicly retracted this support in Norway's major national business daily *Dagens Næringsliv* the next day. In this section, we show that this retraction—unreported by Ahern and Dittmar (2012)—has important implications for the interpretation of their five-day abnormal return estimate. While they assume that their five-day abnormal return estimate reflects the pro-quota Friday announcement only, the daily abnormal return estimates in Panel A of Table 7 show that this assumption is wrong. Again using their estimation methodology, Panel A shows that the market reaction to the Minister's Friday announcement in support of the quota (day 0) is statistically insignificant even *without* correcting their standard errors for cross-sectional dependence.

Table 7 further shows that Ahern and Dittmar (2012)'s negative five-day abnormal return estimate is driven by February 20 (day -2) and February 25 (day +1). Day -2 precedes the Minister's first interview and, hence, does not represent quota-related information. However, since the negative market reaction on day +1 reflects the Minister's Saturday *withdrawal* of his quota support, it logically implies that the market valued the quota *positively.* In other words, Ahern and Dittmar (2012) assign their negative abnormal return to the wrong event, leading to a misinterpretation of their own empirical result.

However, as shown in Panel B of Table 7, none of the event days within the five-day event window are statistically significant after correcting the standard errors for cross-dependence. Thus, our conclusion of a statistically insignificant market reaction to the gender quota is valid also after examining each individual day within Ahern and Dittmar (2012)'s five-day event window.

3 Tobin's Q and operating profitability

In this section, we examine whether the gender quota affected Tobin's Q (henceforth Q) of listed ASA. Under market efficiency, the market reaction to the quota news announcement is an unbiased estimate of the valuation effect. This implies that there should be no systematic changes in market equity value or Q in subsequent periods. However, given uncertainty about firms' ability to appoint qualified female directors, Q may change as firms comply with the quota and investors update their priors. We begin with a reduced-form regression, followed by our implementation of the two-stage instrumental variable (IV) analysis designed by Ahern and Dittmar (2012). As our conclusion differs substantially from theirs, we end the section by carefully identifying the main source of this difference.

3.1 Reduced-form regression

We first estimate the following reduced-form regression, using an unbalanced panel of 239 listed ASA over the period 2002-2008:

$$Q_{it} = \alpha + \beta Shortfall_{it} + \theta_i + \tau_t + \epsilon_{it}, \tag{7}$$

where θ_i and τ_t are firm and year fixed effects, respectively. The data used to compute Q is from *Oslo Børsinformasjon.* Q is defined as (book value of total assets − book value of equity + market value of equity)/(book value of total assets). The market value of equity is the stock price times the number of shares outstanding (shares issued − treasury shares), using the end-of-year closing price. If a firm has more than one share class, the market value of equity is the combined market value of all share classes. We eliminate Q-values that are less than or equal to zero, and winsorize the remaining observations at 1% and 99% each year. The estimation (not tabulated) yields a statistically insignificant coefficient estimate of $\beta = 0.030$, suggesting that the actual shortfall of female directors is unrelated to Q.

3.2 Two-stage IV analysis

We next employ the IV analysis designed by Ahern and Dittmar (2012). By way of motivation, recall first that, while the law required full compliance by the end of 2007, several listed ASA began voluntarily appointing female directors already in 2002, shortly after the first major quota announcement in February of that year. The first-stage in the IV analysis is designed to eliminate effects of firm-specific (endogenous) compliance timing over the six-year compliance period. Essentially, it involves instrumenting the fraction of female directors with the market-wide board trend from a firm-specific starting point at year-end 2001, prior to the first quota news event in early 2002.

In the first stage, we regress $Shortfall_{it}$ on $Shortfall_{iT_0}$ interacted with year dummies (D_t) as follows:

$$Shortfall_{it} = \alpha + \beta_t D_t Shortfall_{iT_0} + \theta_i + D_t + u_{it}, \tag{8}$$

where $Shortfall_{iT_0}$ is firm i's (exogenous) quota-induced shortfall of female directors in base-year T_0. In the second stage, Q is regressed on the predicted shortfall, $\widehat{Shortfall}_{it}$, from the first-stage regression:

$$Q_{it} = \alpha + \beta \widehat{Shortfall}_{it} + \theta_i + \tau_t + \epsilon_{it}. \tag{9}$$

For robustness, because the number of missing female directors is an integer, we also re-estimate the IV test replacing $Shortfall$ with the number of female directors, $Shortfall_{number}$, in the first stage.

The first-stage regression estimates are shown in Panel B of Table 8, while Panel A shows the second-stage estimates. In the first four columns, the estimation period starts in 2002 and ends in 2008, when all the new female directors are in place. Since 2001 is the last year-end in which the cross-sectional distribution of female directors is unaffected by the quota legislative events, $T_0 = 2001$ in Columns (1) and (2). In Column (1) of Panel A, the coefficient estimate for $\widehat{Shortfall}$ is a statistically insignificant 0.750. In Column (2), $\widehat{Shortfall}_{number}$ also receives a statistically insignificant coefficient estimate. In Columns (3) and (4) we move the (exogenous) base-year for $\widehat{Shortfall}$ and $\widehat{Shortfall}_{number}$ back one year to $T_0 = 2000$. Again, there is no evidence of the quota having an impact on Q. While not tabulated, this conclusion is unchanged if we include *Total assets*, *Leverage* (ratio of book value of total debt to total assets), and *Board size* as control variables in the second stage. We conclude from this that Q is unrelated to the shortfall of female directors, also when instrumented with the market-wide time trend.

3.3 A closer look at AD's first-stage instrumentation

While the above analysis closely replicates the main IV test of Ahern and Dittmar (2012), they conclude with a large negative (-20%) and statistically significant impact of the quota on Q. In this section, we explain why our conclusion differs substantially from theirs. This explanation also motivates why our results in the first four columns of Table 8 provide the most appropriate empirical basis for measuring the effect of the quota on Q.

The key to understanding the dramatic difference in inference comes from observing that Ahern and Dittmar (2012) employ year-end 2002 as base-year in their first-stage estimation. Columns (5) and (6) of Table 8 show the effect of repeating the above IV test with $T_0 = 2002$ and with the estimation period 2003 to 2009, which is the same as in Ahern and Dittmar (2012). $\widehat{Shortfall}$ is again replaced by $\widehat{Shortfall}_{number}$ in the even column.

The inference from this exercise is clear: We are able to replicate the second-stage coefficient estimate of Ahern and Dittmar (2012) *only* in Column (5) where $T_0 = 2002$ and the regressor is $\widehat{Shortfall}$. Notwithstanding our larger sample of listed ASA (820 firm-years versus their 603), with this specification, our coefficient estimate is 1.91, which is close to their estimate of -1.94 (Panel A of their Table IV). The sign of our coefficient estimate is switched because we use the shortfall of female directors, while they use the (near-inverse) percent female directors.[7] With the 2003-2009 sample period, the estimated slope coefficient in the second stage is statistically insignificant whether using $\widehat{Shortfall}_{number}$ (Column (6)) or $T_0 = 2001$ (Column (7)). While not tabulated, our main conclusion of a statistically insignificant impact of the quota on Q holds also if we push back the first-stage exogenous base-year to $T_0 = 2000$, keeping the 2003-2009 sample period.

Figure 2 illustrates the impact on the first-stage instrumentation of shifting the base-year from $T_0 = 2001$ to $T_0 = 2002$. Each line in the figure is a firm-specific pathway from 2003 through 2009 using the annual fitted value of $\widehat{Shortfall}$ from the first stage. As explained above, this pathway captures the annual economy-wide trend in the fraction of female directors, with the firm's actual fraction of females in the base-year T_0 as starting point, eliminating the effect of idiosyncratic (endogenous) timing in the hiring of female directors. While the number of unique ASA is 227 in the figure, the number of distinct pathways is much smaller (13 in Panel A and 12 in Panel B of Figure 2). This is because firms with the

[7]While not affecting the results, they use industry-adjusted Q, which we do not (Gormley and Matsa, 2014).

same initial shortfall and board size have the same estimated pathway until 2008.

In Figure 2, the pathways are based on $T_0 = 2001$ in Panel A and on $T_0 = 2002$ in Panel B. The increased cross-sectional dispersion of the pathways from Panel A to Panel B reflect changes in the fraction of female directors from year-end 2001 to year-end 2002. Specifically, a few firms increased their female director shortfall—perhaps by changing board size (top line in the figure), while several others reduced the shortfall to zero or even negative (bottom line in the figure).

There is little doubt that these changes in part reflect the significant increase in the probability of a quota law conveyed by the announcements on February 22 and March 8 of 2002 (Table 2 above). ASA general shareholder meetings (GM) typically take place in April/May, and the meeting agenda with director nominees must be sent out two weeks in advance. Thus, many firms started moving towards compliance with the expected future quota already at the 2002 GM. As discussed above, almost half of the listed ASA reduced their shortfall of female directors between 2001 and 2002.[8] This budding compliance activity invalidates the choice of $T_0 = 2002$ as base-year in the first stage of the IV analysis.

In sum, correcting for the endogeneity of the base-year used by Ahern and Dittmar (2012), there is no evidence of a decline in Q for firms with a more binding quota constraint (i.e., firms with high shortfall). As Tables 8 demonstrates, with a base-year of $T_0 = 2001$ or $T_0 = 2000$ in the instrumentation, there is no evidence that *Shortfall* has an impact on Q in the second stage. Moreover, when using $\widehat{Shortfall}_{number}$ in the second-stage estimation, there is no statistically significant impact on Q even for $T_0 = 2002$. In sum, our Q-evidence corroborates our statistically insignificant abnormal return estimates in Section 2 above.

4 Changes in operating profitability

By adding new inexperienced female directors, the quota could conceivably affect operating profitability (ROA) negatively. Matsa and Miller (2013) report evidence consistent with a negative ROA effect. Since we conclude above that the quota had a neutral impact on equity values, we examine in this section whether the ROA effect is sufficiently large to challenge this conclusion. While Matsa and Miller (2013) use a difference-in-difference test over the period 2003-2009, where the control firms are either Norwegian

[8] Of the 146 listed ASA with board information in both years 2001 and 2002, 13% increased the number of female directors, while 29% reduced the fraction of female director shortfall by reducing board size. We return to board changes in Section 6 below.

AS or listed firms in other Nordic countries, we report difference-in-difference regressions where the 1% largest AS are control firms.

4.1 Sample of treated and control firms

In this section, treated firms include both listed and unlisted ASA. We select as control group so-called "Large AS", defined as the 1% largest firms by revenue among the population of AS. We require AS to have complete accounting information, and exclude subsidiaries not reporting consolidated accounts, using accounting data from the *Brønnøysund Register Centre*. We require AS to have positive total assets and revenue, non-negative long-term assets, current assets, long-term debt, and short-term debt. We also require current assets≥cash and total assets≥working capital (defined as current assets−current debt). This selection results in a total sample of 1150 ASA (listed and unlisted) and 3532 Large AS.

Table 9 reports bi-annual firm and board characteristics for ASA and Large AS over the period 1998-2013. In an average year, there are 174 listed ASA (Panel A), 255 unlisted ASA (Panel B), and 987 Large AS (Panel C). As shown in Columns (1) and (2), the average listed ASA is about five times larger than the average unlisted ASA by revenue and three times larger by total assets. Relative to Large AS, listed ASA have on average twice the revenue and three times the total assets. The ownership data in Column (3) is from the Norwegian tax authorities for 2004-2013 and *Brønnøysund Register Centre* for 1998-2003. Stock ownership is highly concentrated across all three groups. As shown in Column (3), the largest blockholder has an average stake of 34% in listed ASA, 57% in unlisted ASA, and 66% in Large AS.

The quota appears to have had only a negligible spill-over effect on the gender representation of Large AS boards. Column (4) of Table 9 reports the percent female directors on the board. While it increases to 41% by 2009 for ASA, the increase for unregulated Large AS is a modest five percentage points, from 8% in 2001 to 13% in 2013. This modest increase is, however, similar to the general trend throughout western economies. For example, by 2013, the fraction of female directors was on average 18% in EU large publicly traded firms and 17% in US Fortune 1000 firms.[9]

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

4.2 ROA test results

We test whether the quota-induced change in operating profitability for ASA (regulated) differs significantly from that of Large AS (unregulated) by estimating the following OLS panel regression:

$$ROA_{it} = \gamma_0 + \gamma_1 ASA_i * Comply_t + \gamma_2 \mathbf{X_{i,t}} + \theta_i + \tau_t + \epsilon_{it}, \quad (10)$$

where θ_i and τ_t are firm and year fixed effects, respectively. The dependent variable is firm i's ROA in year t, defined as the ratio of earnings before interest and tax (EBIT) to total assets. The indicator variable $Comply_t = 1$ for years $t \geq 2008$, as compliance was required by 12/2007, and $Comply_t = 0$ otherwise. We exclude financial firms and firms switching between the ASA and AS corporate form, and require firms to have at least two observations during the sample period. The sample consists of 559 unique ASA and 1865 unique Large AS, 2001-2013, for a total of 13,779 firm-years.

The vector $\mathbf{X}_{it}$ contains the following firm characteristics (defined above and in Table 4): *Board size*, *Board busyness* (the fraction of directors that hold three or more board seats in ASA or Large AS), *Firm age* (natural log of number of years since incorporation), *Size* (natural log of revenue), *Leverage*, and *Largest owner*. Finally, *Board CEO experience* is the fraction of directors with large-firm CEO experience. *Large-firm CEO experience* is a binary variable with a value of one if a director of firm i is a current outside CEO of an $\text{ASA}_{j \neq i}$ or Large $\text{AS}_{j \neq i}$, or was CEO of any ASA or Large AS in at least one of the last three years. We exclude current inside CEOs since they are prohibited by law to sit on ASA boards after 2010. Since inside CEOs regularly participate in board meetings, boards benefit from the experience of the inside CEO whether or not he/she is a formal board member.

The coefficient estimates are shown in Table 10. In Columns (1) and (2), which use the full sample period 2001-2013, the interaction variable $ASA * Comply$ is statistically insignificant whether the regression controls for firm-fixed effects only (Column 1) or adds firm-specific control variables (Column 2). Thus, there is no evidence that the forced gender balancing of boards affected operating profitability.

Since this conclusion differs from that of Matsa and Miller (2013), Columns (3) and (4) repeat the regressions on the shorter sample period of their paper (2003-2009). Consistent with their finding, the coefficient estimates for $ASA * Comply$ are now negative and significant at the 5% level. The contribution of Table 10 is therefore to show that the negative operating performance of regulated ASA relative to unregulated Large AS is at best short-term—restricted to the recession years 2008 and 2009. That is, for a post-quota period beyond the financial crisis, there is no evidence of the quota law affecting firms' operating performance.[10]

5 Changes in board CEO experience

For the average board, Ahern and Dittmar (2012) report a decline in board CEO experience from 65% of directors in 2002 to 58% in 2009 (their Table II). Moreover, the conclude that "There is a meaningfully large difference in CEO experience of female versus male directors, whether the director is new, retained, or exiting." (p.169). In this section, we construct our own measure of board CEO experience, which is based on our population data on ASA and AS, and we use this measure to show that the decline in board CEO experience documented by Ahern and Dittmar (2012) is unlikely to affect firm value.

Note first that a director's CEO experience in a large firm is likely of greater value than s small-firm experience. Since the population of Norwegian AS is overwhelmingly dominated by tiny firms (two employees or less), we therefore construct two measures of board CEO experience: *Large-firm CEO experience* (defined above), which is restricted to ASA and Large AS, and *Small-firm CEO experience*, which records CEO experience in ASA and *all* AS. We show below that Ahern and Dittmar (2012)'s definition of board CEO experience is dominated by experience in tiny firms. Moreover, when controlling for changes in board CEO experience of Large AS (using Eq. (10) above), we show that ASA board large-firm CEO experience does *not* decline significantly following quota compliance.

5.1 Board large-firm CEO experience

Recall that large-firm CEO experience measures experience in ASA and Large AS with a look-back period of three years and excluding current inside CEO positions. As shown in Figure 3, 8% of ASA female board seats over the period 2001-2013 are held by directors with large-firm CEO experience, versus 19% of male board seats. Also, since, according to Column (6) of Table 9, only 5% of ASA and Large AS CEOs are female, a greater fraction of female CEOs have board seats than do male CEOs. Figure 3 also plots the evolution of the average ASA board large-firm CEO experience, i.e., the fraction of directors with CEO experience averaged across ASA boards. As shown, this measure of board CEO experience changes from

[10]Referring to our ROA evidence, Miller (2018) suggests a labor cost channel: "[the short-term negative effect on ROA documented by Eckbo et al.] is consistent with profits at affected firms being lower during the recession years when they bore additional labor costs from retaining workers, but then rebounding relative to other firms during the recovery."

19% in 2001, 15% in 2008 and 14% in 2013. Column (7) of Table 9 shows this board CEO experience separately for listed and unlisted ASA and Large AS. In 2001, 20% of the directors of the average listed ASA has large-firm CEO experience, changing to 18% in 2009. Notice that these percentages are much lower than the percent of directors with CEO experience reported by Ahern and Dittmar (2012).[11]

We use Eq. (10) to test whether the average change in large-firm CEO experience for ASA boards is significantly different from that of Large AS. The dependent variable is now the fraction of the firm's directors with large-firm CEO experience, and $\mathbf{X_{it}}$ contains *Firm age*, *Size*, *Leverage*, and *Largest owner*. As in Table 10, we use *Size* as a proxy for firm size. Replacing *Size* with *Total Assets* does not change our inferences. As shown in Columns (1) and (2) of Table 11, the estimates of the coefficient γ_1 on $ASA * Comply$ are statistically insignificant, whether or not firm controls are included. These estimates fail to support the hypothesis that ASA-board large-firm CEO experience falls relative to that of Large AS after quota compliance. Moreover, they are consistent with a neutral valuation effect of the quota.

5.2 Board small-firm CEO experience

Ahern and Dittmar (2012) report a statistically significant reduction in CEO experience of listed ASA facing the largest quota constraints. They define CEO experience as "work experience as CEO or owner" based on director bios in company annual reports. Since their definition does not restrict the size of the firm or separates owners from CEOs, it results in a much higher level of board CEO experience than what we report in Column (7) of Table 9. Indeed, with their definition, in an average year over the 2001-2009 period, 66% of male directors and 40% of female directors in their sample have CEO experience (online Appendix Table II).

In order to understand the high percent of female directors with CEO experience, we change our definition of CEO experience to include positions in the remaining 99% of the annual population of roughly 100,000 AS. The results are shown in Column (8) of Table 9. Since 46% of all AS have at most one employee, 58% have at most two employees, and 90% have at most ten employees, we label Column (8) "small-firm CEO experience". The small-firm inclusion raises the average proportion of listed ASA boards' CEO experience in 2001 to 53%, a proportion that is closer to the CEO experience reported by Ahern and Dittmar (2012). Half of the board's CEO experience in listed ASA is now from firms with

[11] While this degree of board CEO experience may seem low, Fahlenbrach, Low, and Stulz (2010) report that only 10% of outside directors in listed US firms are current outside CEOs.

zero to two employees and two-thirds is from firms with at most ten employees. Column (8) shows that the proportion of small-firm CEO experience of the average listed ASA board declines somewhat from 56% in 2001 to 54% in 2013.

In Columns (3) and (4) of Table 11, we use our diff-in-diff regression (Eq. 10) to test whether this decline is statistically significant relative to the change in small-firm CEO experience of Large AS boards. Consistent with the finding of Ahern and Dittmar (2012), the coefficient estimate for $ASA * Comply$ is negative and statistically significant in both regressions. Our contribution here is to show that the decline in ASA board CEO experience is coming from a decline in small-firm CEO experience and *not* from CEO experience in ASA and Large AS. This refinement is important because it suggests that the overall decline in CEO experience reported by Ahern and Dittmar (2012) may be economically unimportant, which is consistent with our valuation analysis.

6 Other board changes

The above evidence shows that, while the fraction of female directors with large-firm CEO experience is generally low, and did not change much over the sample period, boards managed to maintain such large-firm CEO experience. Boards may have achieved this by adjusting board size to make room for more male directors, by not reappointing relatively inexperienced male directors, or by targeting a few but highly qualified females. In this section, we first examine whether ASA made changes to their board size consistent with such adjustments. We then present evidence on board seat concentration and test for changes in board busyness. The idea is that, if qualified females were in short supply or the search cost for such female directors was high, the quota would likely result in a high board seat concentration among a few well-known but "busy" female directors. We end the section by estimating the propensity of listed and unlisted ASA to convert to the unregulated AS legal form as a result of the quota constraint.

6.1 Board size

Recall from Figure 1 that the average board size remains largely unchanged over the sample period. Columns (5) and (6) of Table 11 test whether the average ASA board size differ significantly from those of Large AS throughout the sample period. The table uses regression Eq. (10), but now with board size as dependent variable, 2001-2013. In column (5), the coefficient on the interaction variable $ASA*Comply$

is negative and significant at the 5% level, with $\gamma_1 = -0.184$. However, when adding firm characteristics in Column (6), the significance drops to the 10% level. Thus, there is at best weak evidence that ASA board size declines on average in response to the quota.

Having said that, the board-size frequency distributions in Figure 4, which compares ASA in the two years 2001 (before the start of the public quota debate) and 2008 (when all ASA were in full compliance), show a shift in board size from four and six members to five. In 2001, 37% of the sample firms had boards with four or six directors, which drops to 19% in 2008. Over the same period, the sample proportion of five-member boards increased from 26% to 43%. As noted in Section 2.1 above, this convergence towards a five-member board suggests that there were some movement in board size motivated by the quota. However, the regressions in Table 11 confirms that these changes did not cause board size to increase, as would be the case if shareholders viewed the quota constraint as low cost.

6.2 Board busyness

The valuation effect of exogenous changes in board busyness is an open empirical question. Some studies find positive effects of increased busyness on board decisions and firm value, citing higher director quality and better advice (Masulis and Mobbs, 2011; Field, Lowry, and Mkrtchyan, 2013). Other studies emphasize the potential for negative effects related to lack of time and attention devoted to the firm (Core, Holthausen, and Larcker, 1999; Fich and Shivdasani, 2006; Hauser, 2018).

In Panel A of Figure 5, we plot the frequency distribution of the number of board seats in ASA and Large AS held by male and female directors of 555 ASA in 2001 and 395 ASA in 2008. Two observations stand out. First, ASA directorships are highly dispersed—almost three quarters of individual directors hold only one board seat. Second, following quota compliance (year-end 2007), seat dispersion is largely similar for male and female directors. Using the literature's definition of a busy director—holding three or more board seats (here in ASA and Large AS)—the proportion of busy directors increases from 9% in 2001 to 11% in 2008 for female directors, and from 12% to 13% for male directors.

While Panel A focuses on the number of board seats per individual director, Panel B plots the fraction of busy ASA board directors. The numbers in Panel B are generally higher than those in Panel A because an individual busy director is now counted as many times as he/she has board seats in ASA. Interestingly, the fraction of board seats held by busy female ASA directors increases substantially from 14% in 2001 to 25% in 2008, and drops back to 21% in 2013. For male board seats, it drops from 25% to 20% in 2008 and rebounds somewhat to 22% in 2013. As a result, average board busyness declines slightly following quota compliance.

The decrease in board busyness is confirmed by the regression estimates in Columns (7) and (8) of Table 11. The coefficient estimate for the interaction variable $ASA * Comply$ is negative and significant in both columns. In sum, while female director busyness increased to the prior level of male directors, overall ASA board busyness decreases post-quota relative to unregulated Large AS. Combined with the evidence of no increase in board seat concentration, and the neutral effect on operating profitability and equity value, it suggests that the pool of qualified female directors was sufficiently large to avoid significant adverse economic effects of the quota.

6.3 Conversions from ASA to AS

As first noted by Nygaard (2011), the quota constraint may have pushed some ASA to convert to the unregulated AS corporate form. While both Ahern and Dittmar (2012) and Bøhren and Staubo (2014) argue that a significant number of ASA were in fact pushed by the quota to make this conversion, we demonstrate in this section that the evidence says otherwise. Note at the outset that ASA to AS conversions (and the opposite) occur both before and after the quota compliance period. Moreover, firms generally do not reveal the reason for the conversion. Therefore, inferences as to the true motivation behind a conversion, of which avoiding the quota is one hypothesis, are necessarily indirect. We follow Bøhren and Staubo (2014) and predict that, for conversions to be driven by the quota, the conversion likelihood must increase in the female director shortfall.

We also stress that there are few if any benefits from the ASA corporate form other than to raise public equity (only ASA may list on OSE). Therefore, a straightforward alternative hypothesis is that a converting ASA has abandoned plans to go public. Indeed, for such firms, remaining an ASA involves certain costs, including having to register the shares with Norway's securities registry (VPS), as well as some additional corporate governance restrictions (Bøhren and Staubo, 2014). For example, OSE-listed firms must report their financial statements according to International Financial Reporting Standards (IFRS), comply with or explain any deviation from the OSE corporate governance code, and disclose the compensation of executives and directors. Investors must flag when they cross certain share ownership thresholds, and reaching a 33% ownership fraction triggers a mandatory bid for the remaining shares. Under this alternative hypothesis, the female director shortfall is not a determinant of conversion likelihood.

6.3.1 Listed ASA

Our data sources for conversions are the *Brønnøysund Register Centre*, supplemented with news and press releases to identify cases where an ASA was acquired by a foreign company. We exclude financial ASA because they were required to be incorporated as ASA until 2007. Column (1) of Table 12 lists the beginning-of-year population of non-financial ASA over the period 2002-2009. It shows in Panel A that the number of OSE-listed ASA increased each year from 2004 through 2008. Column (2) lists the total number of firms that exit the ASA legal form each year. Column (3) shows the exits that are due to M&A and bankruptcy. We follow Bøhren and Staubo (2014) and exclude these events from the conversion analysis below since they are much too complex to be driven by quota avoidance (while switching to AS only requires a bylaw change). This leaves the conversions for "other" reasons in Column (4). The research question is whether these conversions are initiated in response to the gender quota. For listed ASA, the answer is immediate: not a single OSE-listed ASA delisted for reasons other than M&A and bankruptcy. Whatever the perceived cost of the gender quota, it was evidently lower than the benefits of remaining publicly listed.

6.3.2 Unlisted ASA

In Panel B of Table 12, Column (4) shows a steady number of conversions by unlisted ASA for reasons other than M&A and bankruptcy. These may be quota-driven or reflecting other reasons, including aborted plans to raise public equity (which may render the less expensive AS form optimal). Column (4) also shows that the number of such conversions by unlisted ASA is higher in 2006 and 2007 than in most other years (except 2002), suggesting that these could be quota-related. To formally test this hypothesis, we estimate the following logit model for the conversion likelihood with the female director shortfall as explanatory variable:

$$Convert_{it} = \alpha + \gamma_1 Shortfall_{it} + \gamma_2 \mathbf{X}_{it} + \kappa_i + \tau_t + \epsilon_{it}, \tag{11}$$

where $Convert_{it}$ takes a value of one if the firm converts to AS in year $t+1$, and zero otherwise. The vector $\mathbf{X}_{it}$ of explanatory variables includes *Board size*, *Firm age*, *Total assets*, *ROA*, *Leverage* and

Largest owner. Firms that convert drop out of the sample in the year of conversion. The unbalanced panel contains 880 firm-years for 264 unlisted non-financial ASA, of which 150 convert sometime in the period 2002-2009.

The results of this logit estimation are in Columns (1)-(4) of Table 13. In the first two columns, the coefficient on *Shortfall* is statistically insignificant. In Columns (3) and (4), $Shortfall$ is replaced with the shortfall number of female directors ($Shortfall_{Number}$), which also receives a statistically insignificant coefficient estimate. This evidence fails to support the hypothesis that that conversions are quota-driven. It does not reject, however, other reasonable alternative explanations, including the possibility that unlisted ASA choose conversion to the lower-cost AS legal form after aborting plans to raise public equity.

While the results in Columns (1)-(4) of Table 13 are new to the quota debate, Bøhren and Staubo (2014) estimate a different conversion likelihood and find that it is increasing in the quota-induced female director shortfall. Most important, they use a different dependent variable, which we label $Convert(BS)_i$. This dependent variable equals one *in all years* for firms that convert. However, because firms drop out of the sample after conversion, the combination of backfilling the dependent variable and the time trend in female director representation produces (mechanically) a higher average $Shortfall$ for converting firms. We control for this time trend by adding year fixed effects. Columns (5)-(8) demonstrate that, when we use $Convert(BS)_i$ as dependent variable in Eq. (11), controlling for the time trend (Columns 5 and 7) eliminates the statistical significance of both $Shortfall$ and $Shortfall_{Number}$. Thus, our conclusion that ASA to AS conversions are unrelated to the quota constraint remains valid.[12]

7 Summary and conclusion

Norway's pioneering board gender quota provides a uniquely important and interesting quasi-experiment for identifying market valuation effects of exogenous changes in director characteristics. On the one hand, prior research on this quota has shown that publicly traded ASA appointed new female directors that were more highly educated professionals than the departing male directors, and that the level of board independence increased, both of which are often associated with more effective boards. On the other hand, prior research also reports that the level of board CEO experience decreased significantly, which

[12]Replacing the variable *Shortfall* with *High shortfall* does not change this conclusion.

may have caused a reduction in board effectiveness and therefore a reduction in equity values.

In this paper, we provide a novel, unified empirical analysis of quota-induced changes in board characteristics and their likely net effect on equity values. Our evidence challenges in particular the widely quoted conclusion in Ahern and Dittmar (2012) that equity values of all-male boards declined substantially due to the quota constraint. We make two simple but important corrections to Ahern and Dittmar (2012)'s econometric methodology. First, we compute standard errors of abnormal returns that account for the substantial cross-dependence in returns that exists when the news event affects all sample firms simultaneously in calendar time. Second, we adjust Ahern and Dittmar (2012)'s instrumented quota-induced annual shortfall in female directors to start in 2001, so that it is truly exogenous to their important quota news announcement in February of 2002.

These corrections are sufficient to render both the quota-induced abnormal stock returns and changes in Q economically and statistically insignificant—whether using Ahern and Dittmar (2012)'s data or our larger set of firms. We also show that, in the cross-section of listed ASA, the valuation estimates do not depend on the quota-induced female director shortfall. In sum, our new valuation estimates, which effectively replace those in Ahern and Dittmar (2012), show that the quota constraint did not materially affect equity values.

We further show that Ahern and Dittmar (2012)'s evidence of a significant decline in board CEO experience is unlikely to affect the quality of the boards' advisory function (and therefore equity values) because it is restricted to director CEO experience in *tiny* firms (zero to two employees). We demonstrate, using nearly the population of ASA and AS, that there is no decline in board *large-firm* CEO experience. Our measure of large-firm CEO experience covers CEO positions in ASA and the 1% largest AS by revenue (Large AS), with a look-back period of three years and excluding current inside CEO positions. Our diff-in-diff analysis, which contrasts ASA (treated firms) with Large AS (unregulated by the quota). Consistent with our conclusion of a value-neutral impact of the quota, the relative drop in small-firm CEO experience for ASA boards is likely irrelevant for the quality of ASA directors' advisory role.

We also corroborate our finding of a value-neutral effect of the quota by clarifying Matsa and Miller (2013)'s evidence of negative operating performance (ROA) after the quota, and by new evidence on board characteristics of relevance for judging the supply of qualified female directors. We show that Matsa and Miller (2013)'s evidence holds only when the regression analysis uses a post-quota period ending in 2009. Our diff-in-diff analysis yields a statistically insignificant change in ROA when extending the time series through 2013. Moreover, the average (five-director) board chose not to increase board size, which would have allowed it to make room for new females without having to let go of male directors. This decision suggests that the cost of expanding board size places an upper bound on the expected cost of quota compliance. Also, consistent with a sufficiently large supply of qualified female directors, neither board seat concentration nor director busyness increased over the quota compliance period.

Finally, we address the claim of Bøhren and Staubo (2014) and others that a lower fraction of female directors is associated with a greater likelihood that an ASA converts to the unregulated AS legal form. We show that this conclusion is driven by a time-trend in unlisted ASA created mechanically by the way Bøhren and Staubo (2014) estimate their conversion likelihood. Correcting for the time-trend (using year fixed effects), which is necessary to properly identify a cross-sectional effect of the quota, eliminates the statistical significance of the female director shortfall. We also show that not a single *listed* ASA converted to AS between 2002 and 2009, and that the number of listed ASA itself increased over that period.

Overall, our findings strongly suggest that neither investors nor firms viewed forced gender-balancing as a particularly costly regulation in economic terms. An interesting avenue for future research is whether this conclusion also holds for the gender quotas subsequently adopted by other western European countries, with their historically different cultural approaches to gender balancing.

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Figure 1
Norwegian ASA board size and proportion of female directors, 1998-2013

The figure shows the average board size (left axis), defined as the number of shareholder-elected directors, and the number (left axis) and fraction (right axis) of female directors. The two vertical lines bracket the two-year quota compliance period (12/2005–12/2007). The sample is 1150 Norwegian ASA, 1998-2013. Board data are from *Brønnøysund Register Centre*.

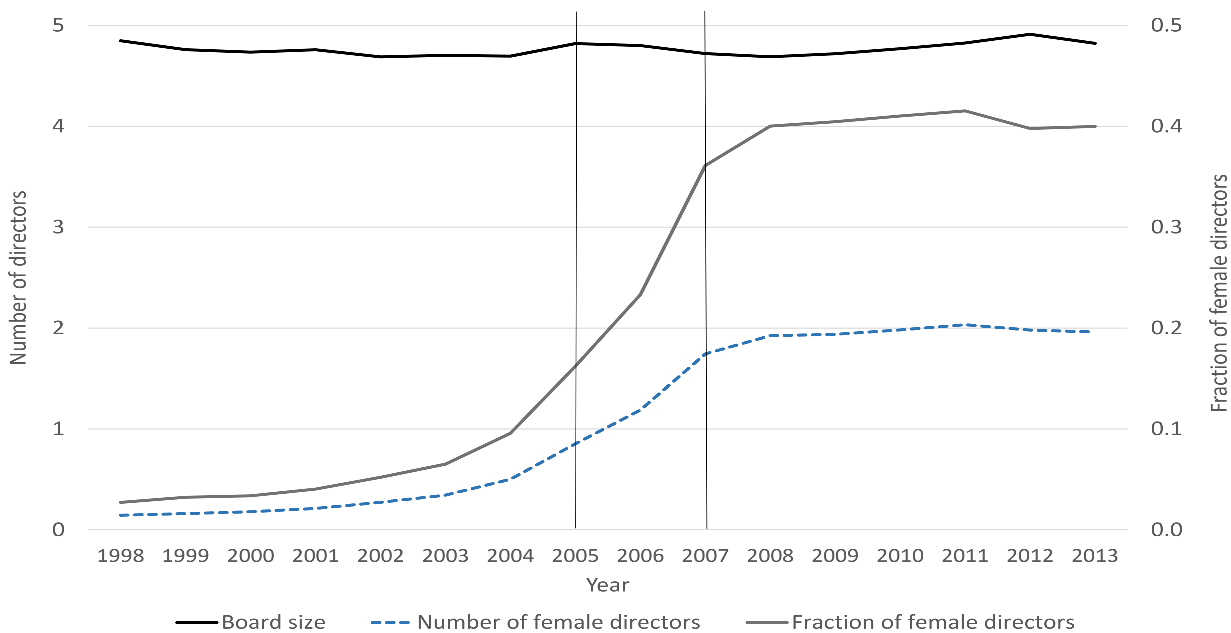


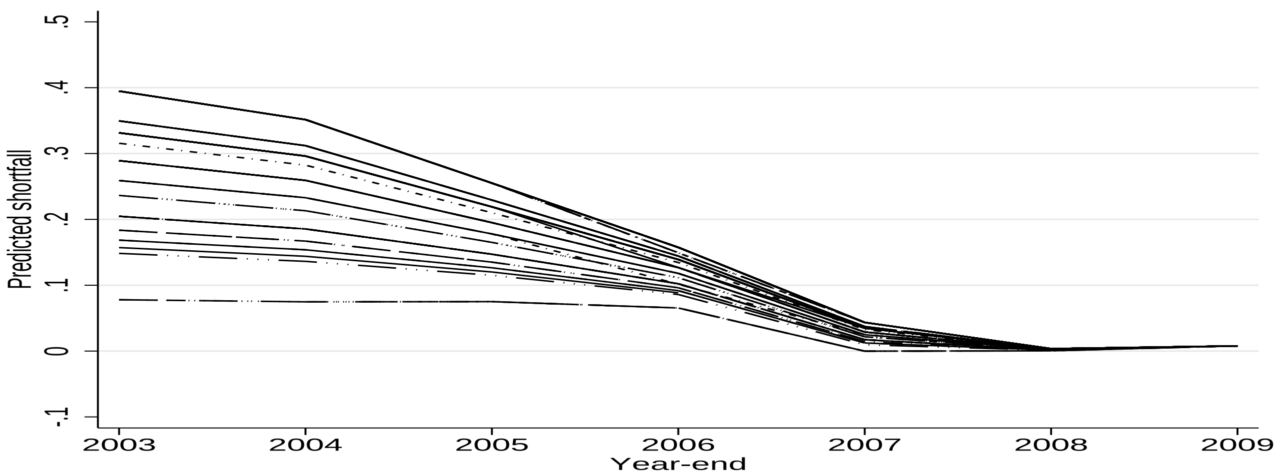
Figure 2
First-stage instrumentation of female director shortfall

The figure plots $\widehat{Shortfall}_{it}$, the predicted shortfall of female directors relative to the quota requirement, estimated from the first-stage regression:

$$Shortfall_{it} = \alpha + \beta_t D_t Shortfall_{iT_0} + \theta_i + D_t + u_{it},$$

where D_t and θ_i are year dummies and firm fixed effects, respectively. $Shortfall_{iT_0}$ is the female director shortfall exogenous to the quota measured at the end of base-year $T_0 = 2001$ (Panel A) and $T_0 = 2002$ (Panel B). Each estimated line moves firm i's initial (exogenous) $Shortfall_{iT_0}$ towards zero (full compliance) at the speed of the sample-wide average change. The sample is 227 listed ASA, 2003-2009.

A: Predicted shortfall $\widehat{Shortfall}_{it}$, with $Shortfall_{iT_0}$ measured in base-year $T_0 = 2001$



B: Predicted shortfall $\widehat{Shortfall}_{it}$, with $Shortfall_{iT_0}$ measured in base-year $T_0 = 2002$

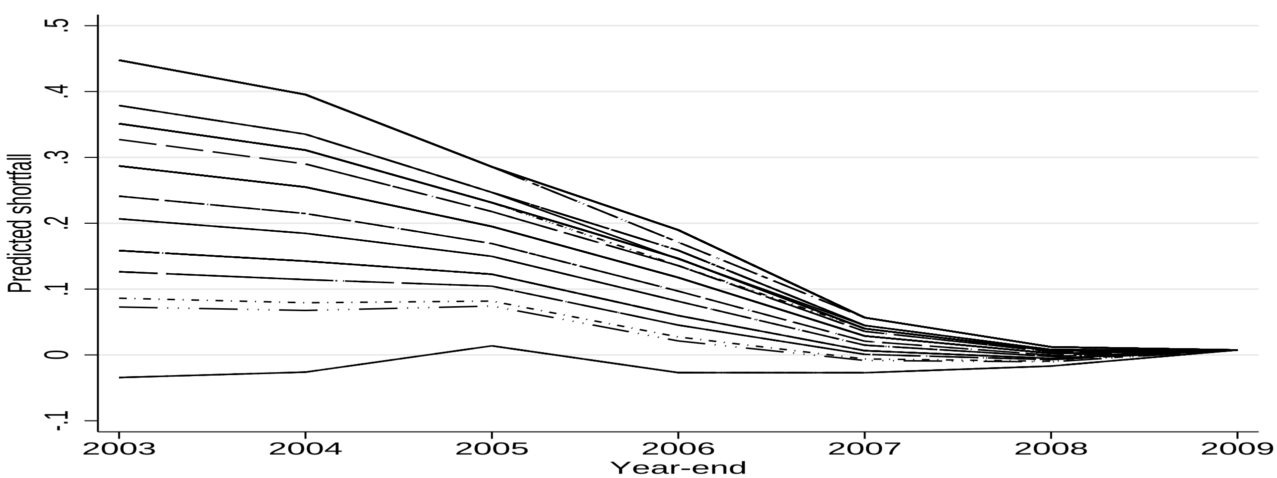


Figure 3
Large-firm CEO experience of ASA directors and boards, 2001-2013

The figure shows the annual fraction of male and female ASA board seats with large-firm CEO experience, as well as the fraction of directors with large-firm CEO experience of the average ASA board. A director of firm i has large-firm CEO experience if he/she is an outside CEO in an $\text{ASA}_{j \neq i}$ or Large $\text{AS}_{j \neq i}$ (top 1% AS by revenue) in year t, or was CEO in any ASA or Large AS in at least one of the past three years. The sample is 5675 male and 1437 female directors of 1150 ASA, 1998-2013. Board data are from *Brønnøysund Register Centre*.

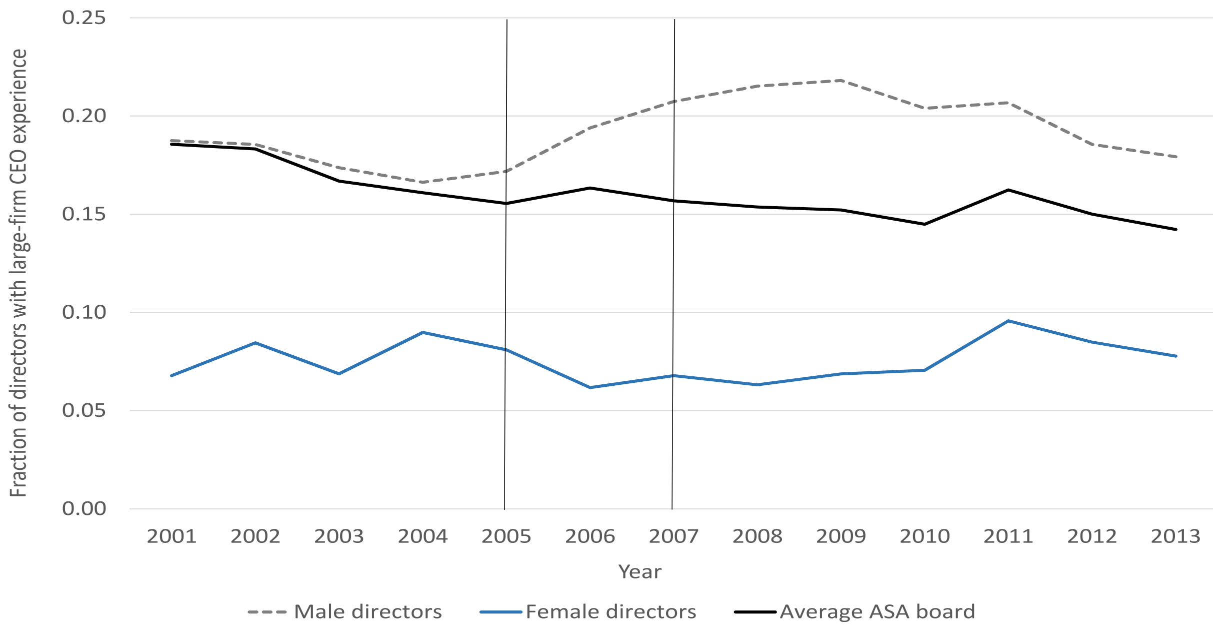


Figure 4
Frequency distribution of ASA board size, 2001 and 2008

The figure shows the frequency distribution of board size. The sample is 555 ASA in 2001 and 395 ASA in 2008. Board data are from *Brønnøysund Register Centre*.

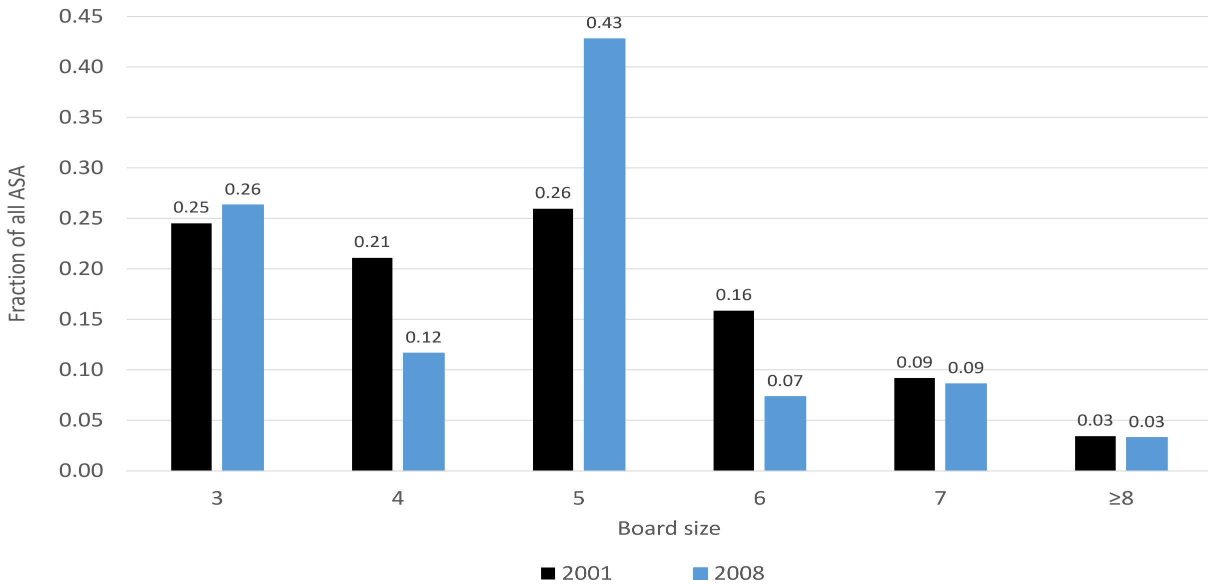
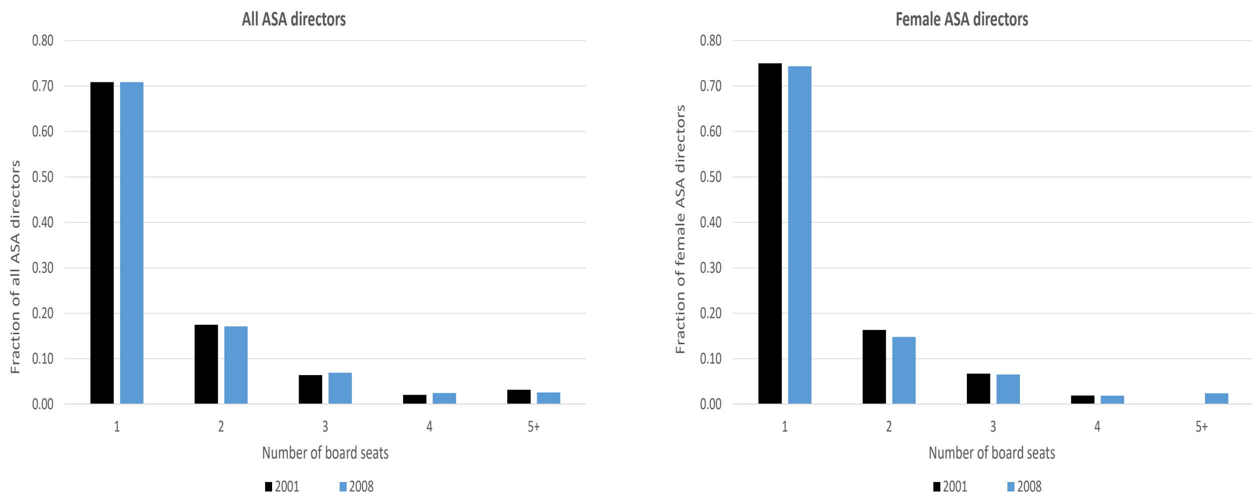


Figure 5
Distribution of ASA director's board seats and average ASA board busyness

Panel A plots the frequency distribution of the total number of board seats in ASA and Large AS (the top 1% AS by revenue) held by ASA directors and female ASA directors in 2001 (prior to quota announcement) and 2008 (after full compliance). Five and more board seats are reported under 5+. The sample is 555 ASA with 2042 directors, of which 104 female, in 2001 and 395 ASA with 1500 directors, of which 581 female, in 2008. Panel B shows the proportion of male and female board seats held by busy directors, as well as the fraction of busy directors on the average ASA board. A busy director holds three or more board seats in ASA and Large AS. The sample is 5675 male and 1437 female directors of 1150 ASA, holding 25,545 and 6276 directorships, respectively, 1998-2013. Board data are from *Brønnøysund Register Centre*.

Panel A: Distribution of ASA directors' board seats in ASA and Large AS, 2001 and 2008



B: Annual fraction of busy ASA directors, 1998-2013

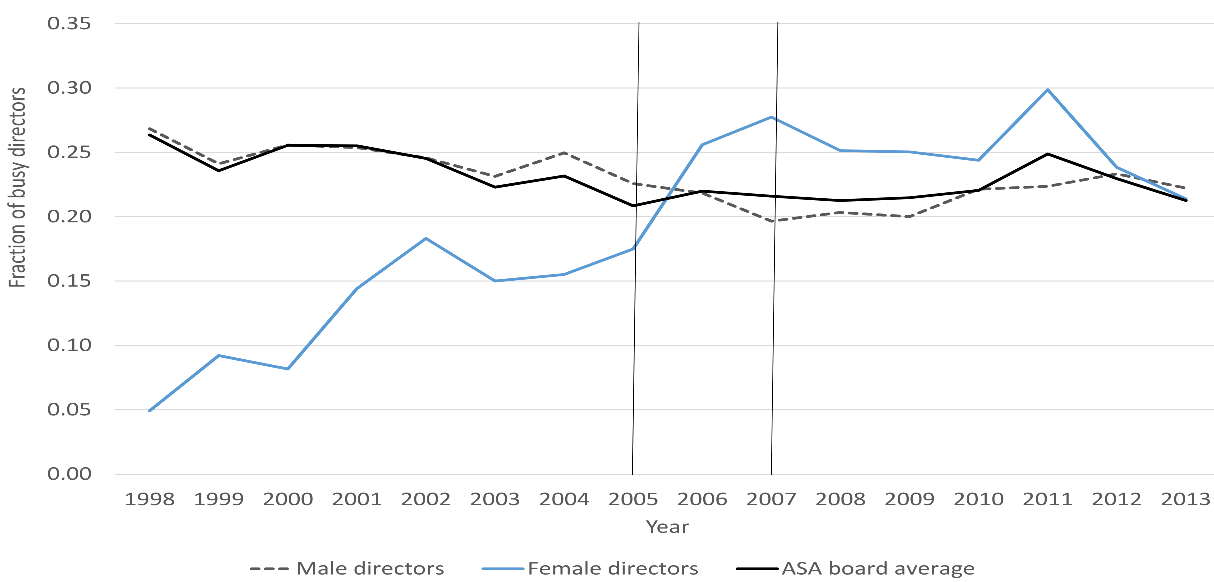


Table 1
Female directors required by Norway's board gender quota

The table shows how the required number and fraction of female directors varies with board size, defined as the number of shareholder-elected directors. $Shortfall$ is the fraction of additional female directors required to comply with the quota for a given board size.

Board size	Required number of female directors (1)	Required fraction of female directors (2)	$Shortfall$ on board with 1 female (3)	2 females (4)	3 females (5)
3	1	0.33	0	0	0
4	2	0.50	0.25	0	0
5	2	0.40	0.20	0	0
6	3	0.50	0.33	0.17	0
7	3	0.43	0.29	0.14	0
8	3	0.38	0.25	0.13	0
9	4	0.44	0.33	0.22	0.11
10	4	0.40	0.30	0.20	0.10
>10	>4	≥0.40			

Table 2
News events increasing the probability of a board gender quota

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(1) **February 22, 2002**: The Minister of Trade and Industry surprisingly supports a gender quota in a newspaper interview, contrary to his political party's official stand. (*Verdens Gang*).

The next day, on Saturday, February 23, 2002, the same Minister retracts his support (*Dagens Næringsliv*). Over the following week, the parliamentary members of his political party publicly reiterates the party's negative stance to a quota.

(2) **March 8, 2002**: In a surprise announcement, Cabinet proposes a board gender quota to be signed into law in 2005. Since the main opposition party already is in favor of a quota, it now has majority support. Cabinet promises compliance by government-owned firms within one year (*Dagens Næringsliv*).

(3) **June 16, 2003**: Cabinet sends the gender quota proposal to Parliament. It contains a sunset provision, cancelling the quota if firms comply voluntarily by 2005. (*Aftenposten*).

During fall, it becomes obvious that the quota has broad political support. As generally expected, the quota proposal passes Parliament's lower chamber (*"Odelstinget"*) on November 27, 2003, and its upper chamber (*"Lagtinget"*) on December 9, 2003. The quota is formally amended to Norwegian Corporate Law on December 19, 2003. To take effect, the quota must be mandated by Cabinet once the sunset provision expires in 2005.

(4) **December 1, 2005**: Female board representation is at 15% and falls short of the quota requirement. The newly elected Prime Minister announces that his Cabinet will mandate the gender quota. There are speculations that the sanction will be a monetary fine (*Verdens Gang*).

(5) **December 9, 2005**: Cabinet mandates the quota, so it takes effect. The ultimate sanction for noncompliance is forced liquidation—the penalty for breach of Corporate Law. Existing ASA are given two years to comply (*Dagens Næringsliv*).

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

The table reports cumulative abnormal stock returns, $CAR_k(-1,0) = 2AR_k$, for portfolios of OSE-listed firms on quota news event k, estimated using the return-generating process:

$$r_t = \alpha + AR_k d_{kt} + \beta_1 w_t + \varepsilon_t,$$

where r_t is the daily equal-weighted return (converted to USD using the daily exchange rate) in excess of the daily 3-month US Treasury bill, d_{kt} is a dummy for the event window (-1,0), and w_t is the daily excess return on the MSCI stock market world index. Events $k = 1, ..., 5$ are defined in Table 2. High shortfall firms have a female director $Shortfall$ (fraction of new female directors required to comply with the quota, see Table 1) at or above the median in the year preceding the event. Portfolios are re-sorted each year-end. Columns (1)–(4) use samples of Norwegian firms, subject to the quota. Columns (5)–(7) use samples of OSE-listed Norwegian (treated) and foreign (control) firms in the oil/offshore sector. N denotes the number of firms in each portfolio. To be included in the portfolio, a firm must have actual return observations both days in the event window and ≥100 return observations in the estimation period (-252, -2). We exclude any earlier event date in the estimation period. For the cumulative return estimate, $CAR_{1-5}(-1,0) = 10AR_{1-5}$, we re-estimate the model with the dummy variable d taking a value of one in the event window (-1,0) for all five events, i.e., a simultaneous estimation of all five event windows. Daily stock returns are from *Oslo Børsinformasjon*. Information on board composition is from *Brønnøysund Register Centre*. Significance levels are *** 1%, ** 5%, * 10%.

	All firms (1)	High shortfall (2)	Low shortfall (3)	High –Low (4)	Domestic oil/offshore (5)	Foreign oil/offshore (6)	Domestic – Foreign (7)
(1) February 22, 2002							
$CAR_1(-1,0) = 2AR_1$	-0.009	-0.012	0.002	-0.013	0.000	-0.019	0.019
p-value	[0.571]	[0.497]	[0.929]	[0.403]	[0.981]	[0.474]	[0.453]
N	143	93	41		32	11	
(2) March 8, 2002							
$CAR_2(-1,0) = 2AR_2$	0.033**	0.036**	0.030	0.007	0.037*	0.051*	-0.013
p-value	[0.035]	[0.037]	[0.102]	[0.682]	[0.052]	[0.053]	[0.602]
N	146	96	41		31	10	
(3) June 16, 2003							
$CAR_3(-1,0) = 2AR_3$	0.000	0.004	-0.006	0.010	-0.015	0.009	-0.023
p-value	[0.989]	[0.831]	[0.771]	[0.585]	[0.549]	[0.811]	[0.512]
N	136	74	53		28	11	
(4) December 1, 2005							
$CAR_4(-1,0) = 2AR_4$	0.001	-0.001	0.003	-0.004	0.009	-0.004	0.012
p-value	[0.941]	[0.946]	[0.811]	[0.538]	[0.643]	[0.887]	[0.467]
N	132	67	65		31	13	
(5) December 9, 2005							
$CAR_5(-1,0) = 2AR_5$	0.009	0.007	0.011	-0.004	0.009	0.016	-0.007
p-value	[0.524]	[0.652]	[0.413]	[0.564]	[0.621]	[0.534]	[0.670]
N	133	67	66		30	13	
All events (1)-(5)							
$CAR_{1-5}(-1,0) = 10AR_{1-5}$	0.038	0.038	0.043	-0.005	0.047	0.062	-0.015
p-value	[0.302]	[0.349]	[0.283]	[0.880]	[0.327]	[0.362]	[0.806]
N	138	79	53		30	12	

Table 4
Definition of variables used in the empirical analysis

The main data source is *Brønnøysund Register Centre*. Data for *Risk* and *Q* are from *Oslo Børsinformasjon*. Ownership data is from the Norwegian tax authorities (2004–2013) and *Brønnøysund Register Centre* (1998–2003). Log refers to the natural logarithm.

Variable name	Definition
A: Firm characteristics	
Firm age	Log of firm age since incorporation.
ROA	Return on assets (earnings before interest and taxes (EBIT) / total assets).
Total assets	Log of book value of total assets.
Size	Log of revenue.
Leverage	Ratio of book value of total debt to total assets.
Largest owner	Percent ownership by the firm's largest shareholder.
Government control	Dummy indicating that the government owns ≥30% of the shares.
Codetermination	Dummy indicating that female directors required by the quota and employee representatives together hold a majority of the board.
Risk	The firm's daily stock return volatility in the year prior to the event.
Q	(Total assets − book value of equity + market value of equity) / total assets.
ASA	Public limited company ("*Allmenaksjeselskap*"). Regulated by the quota.
Large AS	The 1% largest private limited companies ("*Aksjeselskap*" or "AS") by revenue. Not regulated by the quota.
Industry	Firms are allocated to ten different industry sectors: oil/offshore, telecom/technology, manufacturing, construction, wholesale/retail, finance, agriculture, transportation, electricity, other services.
B: Board characteristics	
Board size	Number of shareholder-appointed directors on the board.
Large-firm CEO experience	Dummy indicating that a director of firm i is a current outside CEO of an $\text{ASA}_{j\neq i}$ or Large $\text{AS}_{j\neq i}$, or was CEO of any ASA or Large AS in at least one of the last three years, $t = \{-3, -2, -1\}$.
Small-firm CEO experience	Similar to *Large-firm CEO experience*, but where CEO experience is from any AS, irrespective of size.
Board CEO experience	Fraction of directors with *Large-firm CEO experience*.
Board busyness	Fraction of directors that hold ≥ 3 board seats in ASA or Large AS.
$Shortfall$	Difference between the fraction of female directors required by the quota and that of the current board. See Table 1.
$Shortfall_{Number}$	Difference between the number of female directors required by the quota and that of the current board. See Table 1.
High shortfall	Dummy indicating $Shortfall \geq$ median. In 2007, the median $Shortfall$ is zero and we require $Shortfall > 0$.
Low shortfall	Dummy indicating below median $Shortfall$. *Low shortfall* = *1-High shortfall*.
$Zero_{2001}$	Dummy equal to one if the firm have zero female directors in 2001.
Pos_{2001}	Dummy equal to one if the firm have at least one female director in 2001.
Comply	Dummy equal to one in years $t \geq 2008$ (reflecting quota compliance by 12/2007).

Table 5
Cross-sectional regressions for announcement returns of quota key events

The table reports coefficient estimates β in cross-sectional OLS regressions for the two-day cumulative abnormal return $CAR_{ik}(-1,0)$ on key quota news event dates, $k = 1,\dots,5$ (Table 2). For each firm i, the daily average abnormal return AR_{ik} is estimated for each event k using the regression model in Table 3. The estimation period starts 252 days before each event and ends with the event (day 0). We require firms to have actual return observations on both days in the event window and $\geq$100 return observations in the estimation period. $CAR_{ik}(-1,0) = 2AR_{ik}$ is then regressed on the vector $\mathbf{X}$ of firm characteristics:

$$CAR_{ik}(-1,0) = \alpha_k + \beta_k \mathbf{X_{ik}} + \kappa_i + u_{ik}, \quad k = 1, ..., 5,$$

where $\mathbf{X}$ contains the variables *Shortfall*, *Government control*, *Codetermination*, *Risk*, and *Total assets*, and κ is industry fixed effects. In Panels A and B, $Shortfall$ is the fraction and number, respectively, of missing female directors. All variables are from the year-end prior to the event and defined in Table 4.A constant is included, but not reported. The sample is OSE-listed ASA. Daily stock returns are from *Oslo Børsinformasjon*. Firm and board characteristics are from *Brønnøysund Register Centre*. Robust standard errors (White estimator) are reported in parenthesis. Stars indicate significance levels: *** 1%, ** 5%, * 10%.

	Date of quota news event $k = 1, ..., 5$				
	22-Feb-2002	8-Mar-2002	16-Jun-2003	1-Dec-2005	9-Dec-2005
A: Shortfall measured as fraction of female directors					
$Shortfall$	-0.051	0.017	0.001	-0.013	-0.002
	(0.067)	(0.055)	(0.061)	(0.022)	(0.023)
Firm characteristics:					
Largest owner	-0.001	0.023	-0.021	-0.038*	0.015
	(0.025)	(0.029)	(0.037)	(0.019)	(0.018)
Government control	-0.001	-0.018	0.034	0.010	-0.016*
	(0.025)	(0.021)	(0.026)	(0.009)	(0.009)
Codetermination	0.011	-0.001	0.033**	0.001	0.006
	(0.019)	(0.015)	(0.016)	(0.005)	(0.005)
Risk	-0.773**	0.477**	0.563	-0.473*	-0.290
	(0.310)	(0.236)	(0.504)	(0.284)	(0.717)
Total assets	-0.002	0.009**	0.001	-0.002	0.000
	(0.004)	(0.004)	(0.005)	(0.002)	(0.003)
Industry fixed effects	Yes	Yes	Yes	Yes	Yes
R^2	0.127	0.158	0.152	0.112	0.136
N (firms)	126	128	122	130	131
B: Shortfall measured as number of female directors					
$Shortfall_{number}$	0.004	-0.006	0.002	-0.003	0.000
	(0.011)	(0.009)	(0.008)	(0.004)	(0.004)
Firm characteristics	Yes	Yes	Yes	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes	Yes
R^2	0.124	0.157	0.152	0.116	0.136
N (firms)	126	128	122	130	131

Table 6
Replicating and adjusting Ahern and Dittmar (2012)'s event study

Panel A lists the average cumulative abnormal returns $CAR^{AD}(-2,2)$ reported by Ahern and Dittmar (2012) (AD) for 94 OSE-listed ASA. $CAR^{AD}(-2,2)$ is estimated over a five-day event-window around February 22, 2002 (event 1 in Table 2) as follows:

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

r_i is the return to ASA i and r_{imatch} is the average return to US-listed companies in firm i's GICS industry. Return data are from Compustat Global (Norwegian firms) and CRSP (matched firms). AD report p-values (in square brackets) only, so the corresponding standard error (in parentheses) of the average $CAR^{AD}(-2,2)$ is computed by us. Panel B shows our replication using AD's methodology. Panel C shows the portfolio estimate of the five-day abnormal return using the time series regression in Table 3, but with the equal-weighted industry-adjusted portfolio return, r_{pt}^{-I}, as dependent variable:

$$r_{pt}^{-I} = \alpha + \sum_{\tau=-2}^{2} AR_\tau d_t + \epsilon_t,$$

where $r_{pt}^{-I} \equiv \frac{1}{N}\sum_{i=1}^{N}(r_i - r_{imatch})_t$. Panels B and C use AD's sample and return data sources, and the p-values in square brackets use the standard errors in the line above. We use board data from *Brønnøysund Register Centre*, where 69 firms (vs. 68 in AD) have zero female directors in 2001 ($Zero_{2001}$). $Pos_{2001} = 1 - Zero_{2001}$. Significance levels: *** 1%, ** 5% and * 10%.

	All firms in AD (1)	AD firms with $Zero_{2001}$ (2)	AD firms with Pos_{2001} (3)	Difference $Zero - Pos$ (2) - (3)
A: Original AD CAR estimates in %, no adjustment for cross-dependence of returns				
Average $CAR^{AD}(-2,2)$	-2.573***	-3.547***	-0.024	-3.523***
St.err. of $CAR^{AD}(-2,2)$: $\sigma_{CAR} = \frac{\sigma}{\sqrt{N}}$	(0.757)	(1.030)	(0.824)	(1.297)
p-value	[0.001]	[0.001]	[0.977]	[0.008]
N (firms)	94	68	26	94
B: Replication of AD CAR estimates in %, no adjustment for cross-dependence of returns				
Average $CAR^{AD}(-2,2)$	-2.733***	-3.738***	0.042	-3.780***
St.err. of $CAR^{AD}(-2,2)$: $\sigma_{CAR} = \frac{\sigma}{\sqrt{N}}$	(0.780)	(0.973)	(1.011)	(1.403)
p-value	[0.001]	[0.000]	[0.967]	[0.009]
N (firms)	94	69	25	94
C: Time-series estimation of CAR in %, adjusting for cross-dependence of returns				
Average daily AR over event window (-2,2)	-0.423	-0.661	0.210	-0.873
St.err. of AR: σ_{AR}	(0.650)	(0.685)	(0.711)	(0.563)
Five day $CAR(-2,2) = 5AR$	-2.116	-3.305	1.051	-4.365
Std.error $CAR(-2,2)$: $\sigma_{CAR} = 5\sigma_{AR}$	(3.250)	(3.435)	(3.555)	(2.815)
p-value from $t = \frac{CAR}{\sigma_{CAR}} = \frac{AR}{\sigma_{AR}}$	[0.516]	[0.336]	[0.768]	[0.122]
N (firms)	94	69	25	94

Table 7
Five one-day abnormal returns centered on February, 22, 2002

The table repeats the estimation reported in Table 6 for each day in the five-day event window (-2,2) around February 22, 2002 (event 1 in Table 2). Panel A reports average daily AR for the sample firms in Ahern and Dittmar (2012) (AD), using their methodology:

$$AR_{i\tau} = (r_i - r_{i,match})_\tau$$

where r_i is the return on ASA i on day $\tau = \{-2, -1, 0, 1, 2\}$ in the five-day event window, and $r_{i,match}$ is the average return to US-listed companies in firm i's GICS industry. Return data is from Compustat Global (Norwegian firms) and CRSP (US firms). The sample size changes slightly across event days because some of AD's 94 ASA have missing return observations on some days in the five-day window (-2,2). Panel B shows AR coefficient estimates from the model

$$r_{pt}^{-I} = \alpha + \sum_{\tau=-2}^{2} AR_\tau d_t + \epsilon_t,$$

where $r_{pt}^{-I} \equiv \frac{1}{N}\sum_{i=1}^{N}(r_i - r_{i,match})_t$ is the equal-weighted portfolio of industry-matched returns and d_t is a dummy variable that takes the value of one on event day τ. The estimation period starts one year prior to day $\tau = -2$ and ends on day $\tau = 2$. The p-values (in parentheses) are adjusted for cross-sectional dependence in Panel B, but not in Panel A. Significance levels *** 1%, ** 5% and * 10%.

	All firms in AD (1)	AD firms with 0 $Zero_{2001}$ (2)	AD firms with Pos_{2001} (3)	Difference $Zero - Pos$ (2) - (3)
A: AD daily AR estimates (unadjusted for cross-dependence)				
February 20 (Wednesday)	-1.029**	-1.213*	-0.529	-0.684
	[0.044]	[0.061]	[0.471]	[0.479]
	78	57	21	
February 21 (Thursday)	-0.088	-0.547	1.199**	-1.746**
	[0.866]	[0.418]	[0.030]	[0.042]
	76	56	20	
February 22 (Friday): Announcement	-0.886	-1.028	-0.558	-0.470
	[0.180]	[0.265]	[0.299]	[0.656]
	73	51	22	
February 23 (Saturday): Reversal announcement. OSE closed, no trading.				
February 25 (Monday)	-1.941***	-2.377***	-0.697	-1.680**
	[0.000]	[0.001]	[0.102]	[0.037]
	77	57	20	
February 26 (Tuesday)	0.605	0.562	0.720	-0.158
	[0.189]	[0.341]	[0.271]	[0.855]
	73	53	20	
B: Portfolio estimates of AR (adjusted for cross-dependence)				
February 20 (Wednesday)	-0.785	-0.954	-0.346	-0.610
	[0.588]	[0.533]	[0.828]	[0.627]
February 21 (Thursday)	0.157	-0.288	1.382	-1.672
	[0.914]	[0.851]	[0.385]	[0.184]
February 22 (Friday)	-0.641	-0.768	-0.374	-0.395
	[0.658]	[0.615]	[0.814]	[0.753]
February 25 (Monday)	-1.696	-2.117	-0.514	-1.605
	[0.243]	[0.167]	[0.746]	[0.202]
February 26 (Tuesday)	0.850	0.822	0.903	-0.083
	[0.558]	[0.591]	[0.570]	[0.947]

Table 8
IV regressions for Tobin's Q and female director shortfall

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

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

where θ_i and τ_t are, respectively, firm and year fixed effects. $\widehat{Shortfall}$ is the fitted value from the first-stage IV regression, reported in Panel B:

$$Shortfall_{it} = \alpha + \beta_t D_t Shortfall_{iT_0} + \theta_i + D_t + u_{it},$$

where D_t is a year dummy and $Shortfall_{iT_0}$ is firm i's (exogenous) quota-induced shortfall of female directors (Table 1) in base-year T_0. The table uses base-year T_0 = 2001 (columns 1-2 and 7-8), T_0 = 2000 (Columns (3-4), and T_0 = 2002 (Columns 5-6). $Shortfall$ fraction of female director shortfall) is replaced with $Shortfall_{Number}$ (number of female director shortfall) in the even-numbered columns. All variables are defined in Table 4. The sample is 239 OSE-listed ASA, 2002-2008 (Columns 1-4) and 227 OSE-listed ASA, 2003-2009 (Columns 5-8). Standard errors clustered by firm are reported in parenthesis. Significance levels: *** 1%, ** 5%, and * 10%.

	Sample period: 2002-2008				Sample period: 2003-2009			
	T_0=2001		T_0=2000		T_0=2002 (AD)		T_0=2001	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A: Second-stage IV regression for Q								
$\widehat{Shortfall}$	0.750		0.535		1.910**		0.689	
	(0.737)		(0.757)		(0.833)		(1.236)	
$\widehat{Shortfall}_{Number}$		-0.028		0.043		0.154		-0.149
		(0.139)		(0.144)		(0.124)		(0.242)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
F-statistic	18.53	17.84	19.75	21.34	15.85	15.79	16.69	15.56
N (firm-years)	815	815	726	726	820	820	790	790
B: First-stage IV regression for $Shortfall$								
$Shortfall_{T_0} \times D_{2002}$	0.714***	0.642***	0.743***	0.634***				
	(0.096)	(0.093)	(0.102)	(0.083)				
$Shortfall_{T_0} \times D_{2003}$	0.611***	0.518***	0.655***	0.522***	0.964***	0.911***	0.634***	0.511***
	(0.129)	(0.122)	(0.121)	(0.110)	(0.091)	(0.091)	(0.131)	(0.131)
$Shortfall_{T_0} \times D_{2004}$	0.536***	0.408***	0.637***	0.434***	0.843***	0.704***	0.553***	0.401***
	(0.114)	(0.106)	(0.096)	(0.102)	(0.108)	(0.099)	(0.120)	(0.117)
$Shortfall_{T_0} \times D_{2005}$	0.348***	0.242**	0.439***	0.315***	0.544***	0.475***	0.360***	0.231**
	(0.105)	(0.100)	(0.103)	(0.117)	(0.108)	(0.100)	(0.116)	(0.117)
$Shortfall_{T_0} \times D_{2006}$	0.177**	0.134	0.268***	0.141	0.433***	0.345***	0.184*	0.116
	(0.081)	(0.088)	(0.088)	(0.091)	(0.096)	(0.082)	(0.095)	(0.101)
$Shortfall_{T_0} \times D_{2007}$	0.082*	0.072*	0.046	0.023	0.167**	0.139**	0.087	0.055
	(0.045)	(0.043)	(0.053)	(0.039)	(0.072)	(0.056)	(0.055)	(0.053)
$Shortfall_{T_0} \times D_{2008}$					0.058	0.030*	0.006	-0.010
					(0.043)	(0.017)	(0.044)	(0.030)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
F-statistic	84.79	68.97	76.73	70.15	85.88	63.33	45.55	43.23
N (firm-years)	832	832	740	740	829	829	799	799

Table 9
Firm and board characteristics for ASA and Large AS, 1998-2013

The table reports firm and board characteristics for listed ASA (Panel A), unlisted ASA (Panel B), and Large AS (Panel C), 1998-2013. The mean revenue and total assets are reported in million 2013 USD and winsorized at the 1% tails. Board CEO experience is the fraction of directors with CEO experience in the average firm. In column (7), CEO experience is limited to ASA and Large AS (the top 1% AS by revenue), while in Column (8), it covers all ASA and AS. See Table 4 for definitions and data sources. The last row in each panel lists the pooled average across all firm-years, with the exception of the number of firms, which lists the average annual N over the sample period. The sample is 409 listed ASA, 888 unlisted ASA, and 3532 Large AS.

Year	Number of firms (N)	Mean revenue (1)	Mean total assets (2)	% ownership by largest shareholder (3)	% female directors (4)	% female chairs (5)	% female CEOs (6)	Board CEO experience in large firms (7)	Board CEO experience in small firms (8)
A: Sample of listed ASA									
1998	196	409	756	46	3.3	1.6	2.7	–	–
2001	169	417	827	29	5.3	0.6	3.8	20.5	56.0
2003	151	391	814	30	9.9	2.1	2.1	17.8	53.8
2005	174	446	910	31	22.2	1.7	1.1	17.0	53.5
2007	205	495	1098	33	38.9	2.9	2.9	18.5	53.1
2009	172	549	1127	36	41.1	4.1	3.5	17.7	50.6
2011	172	663	1236	34	41.5	10.5	4.1	17.2	51.7
2013	150	699	1373	34	41.6	8.8	4.7	15.6	54.2
Mean	174	501	1008	34	24.3	3.6	2.8	17.5	53.0
B: Sample of unlisted ASA									
1998	247	46	133	54	2.3	0.4	1.7	–	–
2001	418	56	110	53	3.5	1.3	5.6	17.8	51.4
2003	346	74	226	56	5.0	2.8	5.7	16.2	48.0
2005	275	103	308	59	12.5	2.9	4.0	14.6	47.5
2007	260	93	369	58	33.9	6.9	6.2	13.4	44.0
2009	167	111	586	57	39.8	10.2	9.8	12.7	41.8
2011	126	197	651	55	41.5	11.1	13.6	15.0	46.9
2013	86	229	787	52	37.1	15.1	9.6	11.8	44.3
Mean	255	90	290	57	15.2	4.2	5.3	14.6	46.6
C: Sample of Large AS									
1998	918	174	180	75	7.4	3.3	2.3	–	–
2001	949	120	165	66	8.4	2.5	3.5	12.4	39.8
2003	967	165	193	65	10.1	3.7	3.6	12.7	39.4
2005	887	200	253	64	12.5	4.3	5.3	13.0	43.4
2007	981	273	380	68	13.5	4.6	5.6	15.5	44.8
2009	989	283	445	66	13.4	5.7	5.3	14.2	46.4
2011	1019	313	490	65	13.8	6.2	7.0	13.4	46.3
2013	1158	293	479	65	13.2	5.5	7.2	13.4	47.3
Mean	987	224	319	66	11.6	4.4	4.9	13.7	44.1

Table 10
Quota-induced changes in operating profitability

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

$$ROA_{it} = \gamma_0 + \gamma_1 ASA_i * Comply_t + \gamma_2 \mathbf{X_{it}} + \theta_i + \tau_t + \epsilon_{it},$$

where θ_i and τ_t are firm and year fixed effects, respectively. The dependent variable is firm i's operating profitability (ROA) in year t, defined as earnings before interest and tax (EBIT) / total assets. $Comply_t =$ 1 for year $t \geq 2008$, reflecting mandatory compliance by 12/2007, and zero otherwise. The vector $\mathbf{X}_{it}$ contains the following firm characteristics: *Board size*, *Board CEO experience*, *Board busyness*, *Firm age*, *Size*, *Leverage*, *Largest owner*, and a constant (not reported). All variables are defined in Table 4. The sample comprises 559 ASA (treated firms) and 1865 Large AS (control firms), 2001-2013. We exclude financial firms and Large AS registered as ASA at some point during the sample period, and require firms to have at least two observations in 2001-2013 and all control variables. The sample period is 2001–2013 in Columns (1)–(2) and 2003–2009 in Columns (3)–(4). Standard errors clustered by firm are reported in parenthesis. Stars indicate significance levels: *** 1%, ** 5%, and * 10%.

	2001-2013		2003-2009	
	(1)	(2)	(3)	(4)
$ASA * Comply$	0.004	0.012	–0.025**	–0.023**
	(0.011)	(0.011)	(0.012)	(0.011)
Board size		–0.006***		–0.002
		(0.002)		(0.002)
Board CEO experience		–0.011		–0.015
		(0.010)		(0.014)
Board busyness		–0.006		–0.001
		(0.009)		(0.012)
Firm age		0.024***		0.011
		(0.007)		(0.008)
Size		0.033***		0.036***
		(0.005)		(0.004)
Leverage		–0.243***		–0.264***
		(0.026)		(0.032)
Largest owner		0.026**		0.035***
		(0.013)		(0.012)
Firm fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
R^2	0.024	0.109	0.025	0.136
N (firm-years)	13,779	13,779	7365	7365

Table 11
Quota-induced changes in board characteristics

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

$$Y_{i,t} = \gamma_0 + \gamma_1 ASA_i * Comply_t + \gamma_2 \mathbf{X_{it}} + \theta_i + \tau_t + \epsilon_{it},$$

where θ_i and τ_t are firm and year fixed effects, respectively. The dependent variable Y_{it} is firm i's board CEO experience (columns 1-4), board size (columns 5-6), and board busyness (columns 7-8) in year t. Board large-firm CEO experience is the fraction of firm i's directors that are outside CEOs in $\text{ASA}_{j \neq i}$ or Large $\text{AS}_{j \neq i}$ (the top 1% by revenue) in year t, or were CEO in any ASA or Large AS at some point in the past three years. Board small-firm CEO experience includes directors' CEO experience in *any* AS, regardless of size. Board size is the number of shareholder-elected directors of firm i. Board busyness is the fraction of directors with at least three board seats in ASA and/or Large AS. $Comply_t = 1$ for years $t \geq 2008$, reflecting mandatory compliance by 12/2007, and zero otherwise. The vector $\mathbf{X}_{it}$ contains the following firm characteristics: *Firm age*, *Size*, *Leverage*, and *Largest owner*. All variables are defined in Table 4. A constant is included but not reported. The sample is 559 ASA (treated firms) and 1865 Large AS (control firms), 2001-2013. We exclude financial firms and Large AS registered as ASA at some point during the sample period, and require firms to have at least two observations. Standard errors clustered by firm are reported in parenthesis. Stars indicate significance levels: *** 1%, ** 5%, and * 10%.

	Board large-firm CEO experience		Board small-firm CEO experience		Board size		Board busyness	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
*ASA * Comply*	–0.007	–0.005	–0.048***	–0.047***	–0.184**	–0.129*	–0.033**	–0.034***
	(0.011)	(0.011)	(0.015)	(0.015)	(0.078)	(0.077)	(0.013)	(0.013)
Firm age		–0.019**		-0.011		–0.016		–0.015*
		(0.008)		(0.010)		(0.051)		(0.008)
Size		0.002		–0.002		0.067***		0.009***
		(0.002)		(0.004)		(0.017)		(0.004)
Leverage		–0.022*		0.037**		–0.353***		–0.029
		(0.013)		(0.019)		(0.093)		(0.018)
Largest owner		–0.000		–0.063***		–0.729***		0.033*
		(0.016)		(0.019)		(0.098)		(0.019)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.003	0.005	0.007	0.011	0.004	0.024	0.006	0.009
N (firm-years)	13,779	13,779	13,779	13,779	13,779	13,779	13,779	13,779

Table 12
Exits from ASA and conversions to AS

The table shows annual changes in the sample of non-financial ASA through exit and entry. Column (1) lists the number of firms at the beginning of the year. Column (2) shows the total number of firms exiting the ASA legal form, split by exits due to M&A or bankruptcy (Column 3) and conversions from ASA to AS for all other reasons (Column 4). Column (5) lists the number of entries (new ASA) and Column (6) shows the total number of firms at year-end. The sample is 288 listed (Panel A) and 467 unlisted (Panel B) non-financial ASA, 2002-2009.

Year	# of ASA at beginning of year (1)	All exits (2)	# of exits due to: M&A or bankruptcy (3)	# of exits due to: All other reasons (4)	All entries (5)	# of ASA at year-end (6)
A: Listed ASA						
2002	157	10	10	0	1	148
2003	148	12	12	0	4	140
2004	140	11	11	0	15	144
2005	144	8	8	0	25	161
2006	161	16	16	0	21	166
2007	166	19	19	0	48	195
2008	195	18	18	0	7	184
2009	184	21	21	0	1	162
B: Unlisted ASA						
2002	288	52	25	27	22	258
2003	258	54	35	19	17	221
2004	221	33	14	19	22	210
2005	210	64	45	19	15	161
2006	161	40	15	25	32	153
2007	153	54	27	27	27	126
2008	126	27	15	12	6	105
2009	105	24	16	8	1	70

Table 13
Determinants of the conversion likelihood for unlisted ASA

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

$$Convert_{it} = \alpha + \gamma_1 Shortfall_{it} + \gamma_2 \mathbf{X}_{it} + \kappa_i + \tau_i + \epsilon_{it},$$

where $\kappa_i + \tau_i$ are, respectively, industry and year fixed effects. In Columns (1)–(4), the dependent variable $Convert_{it}$ takes the value of one if the firm converts to AS in the next year, and zero otherwise. In Columns (5)–(8), the dependent variable is $Convert(BS)_i$, which takes the value of one for firms that convert to AS, and zero for firms that do not convert. Firms drop out of the sample when they convert. The explanatory variables are *Shortfall* (columns 1-2 and 5-6) or $Shortfall_{Number}$ (columns 3-4 and 7-8), the control variables in $\mathbf{X}_{it}$ (*Board size*, *Board CEO experience*, *Firm age*, *ROA*, *Total assets*, *Leverage*, and *Largest owner*), industry fixed effects, and a constant (unreported). Year-fixed effects are included in odd-numbered columns only. All variables are defined in Table 4. The sample is 264 non-financial unlisted ASA, 2001-2008, of which 150 convert and 114 do not convert to AS in the period 2002-2009. We exclude firms that exit the ASA legal form due to M&A and bankruptcy (listed in column (3) of Table 12). Standard errors clustered by firm are reported in parenthesis. Significance levels: *** 1%, ** 5%, and * 10%.

	$Convert_{it}$				$Convert(BS)_i$			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$Shortfall$	0.905	0.716			1.729*	4.497***		
	(0.759)	(0.551)			(0.996)	(0.752)		
$Shortfall_{Number}$			0.114	0.088			0.279	0.900***
			(0.160)	(0.120)			(0.190)	(0.162)
Board size	0.063	0.043	0.034	0.024	0.014	0.039	–0.065	–0.214
	(0.082)	(0.078)	(0.101)	(0.094)	(0.118)	(0.115)	(0.139)	(0.145)
Board CEO experience	0.511	0.492	0.525	0.522	0.635	0.673	0.680	0.769
	(0.654)	(0.602)	(0.655)	(0.602)	(0.713)	(0.655)	(0.720)	(0.648)
Firm age	0.078	0.089	0.079	0.093	0.013	0.078	0.014	0.095
	(0.104)	(0.099)	(0.103)	(0.099)	(0.166)	(0.159)	(0.165)	(0.156)
ROA	–0.544*	–0.397	–0.545*	–0.395	–0.273	–0.104	–0.283	–0.074
	(0.282)	(0.276)	(0.283)	(0.277)	(0.418)	(0.415)	(0.408)	(0.399)
Total assets	–0.146**	–0.134**	–0.145**	–0.137**	–0.271***	–0.307***	–0.265***	–0.293***
	(0.062)	(0.057)	(0.061)	(0.057)	(0.103)	(0.102)	(0.100)	(0.097)
Leverage	0.490	0.379	0.505*	0.391	1.380***	1.113***	1.396***	1.116***
	(0.304)	(0.282)	(0.302)	(0.281)	(0.453)	(0.421)	(0.450)	(0.412)
Largest owner	1.605***	1.536***	1.542***	1.491***	1.107*	1.422***	0.990*	1.179**
	(0.377)	(0.350)	(0.369)	(0.347)	(0.581)	(0.576)	(0.571)	(0.557)
Industry fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	No	Yes	No	Yes	No	Yes	No
Pseudo R^2	0.088	0.058	0.087	0.057	0.260	0.211	0.258	0.200
Firm-years	880	880	880	880	880	880	880	880

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