Board Gender-Balancing and Firm Value

Finance Working Paper N° 463/2016
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Abstract

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

JEL Classifications: G38

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BOARD GENDER-BALANCING AND FIRM VALUE*

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Abstract

Studies of Norway's pioneering gender balancing of corporate boards conclude that the gender quota law imposed large costs on shareholders of listed firms. We show that this conclusion does not survive simple econometric adjustments, whether applied to the prior data sets or our substantially expanded panel. With the improved econometrics, our estimates of abnormal stock returns and changes in Tobin's Q fail to reject the hypothesis of a zero impact on firm value. We also document that boards' large-firm CEO experience did *not* decline, which is consistent with a deep pool of qualified female directors.

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"[Norway's gender quota] led to value losses of upwards of 20% for the firms with [all-male boards]...[The] value losses are persistent across time."
— Kenneth R. Ahern and Amy Dittmar, *Quarterly Journal of Economics*, 2012 (abstract).

In 2005, the Norwegian government required public limited companies (ASA) to have gender-balanced boards within two years, with forced liquidation as penalty for non-compliance. Eight other European countries (Belgium, France, Germany, Iceland, Italy, the Netherlands, Portugal, and Spain) and the state of California have since passed their own mandatory board gender quotas, all with lower penalties for non-compliance. As laid out in the Norwegian government's white paper supporting the parliamentary vote (*Odelstingsproposisjon* 97, 2002-2003), the pioneering board quota was driven by gender politics unrelated to firm performance. Thus, its passage provides a rare, quasi-experimental setting for assessing causal effects of exogenous changes in board composition on firm value.

In principle, restricting shareholders' free choice of directors can reduce board effectiveness. For example, firms may be forced to appoint new female directors with less chief executive officer (CEO) experience than male directors, which can reduce firm value (Fich, 2005; Fahlenbrach, Low, and Stulz, 2010). It has also been suggested that relatively independent but inexperienced female directors are overly focused on monitoring and exhibit excessive risk aversion (Adams and Ferreira, 2009).[1] On the other hand, board effectiveness can increase if the gender quota improves otherwise inefficient board elections, perhaps caused by an existing male director "old boys" network (Agarwal, Qian, Reeb, and Sing, 2016). Thus, it is in principle possible for shareholders to benefit from the diversity and broader skill set resulting from adding female directors (Kim and Starks, 2016; Bernile, Bhagwat, and Yonker, 2018; Adams, Akyol, and Verwijmeren, 2018).

To address these and other important hypothetical tradeoffs, we provide new and wide-ranging evidence on the effect of Norway's forced gender balancing on firm value and board characteristics. As it turns out, the empirical evidence that we present fundamentally changes the prior scientific narrative, which hold that the quota imposed economically large and statistically significant costs

[1]See also Linck, Netter, and Yang (2008), Duchin, Matsusaka, and Ozbas (2010) and Masulis and Mobbs (2011) for discussions of the role of director independence. Croson and Gneezy (2009), Adams and Funk (2012) and Sila, Gonzales, and Hagendorff (2016) discuss female director risk aversion.

on shareholders. The most direct support for this prior narrative comes from Ahern and Dittmar (2012) (henceforth AD) who present much-cited evidence of a large decline in equity values (above quote).[2] AD's conclusion receives further indirect support by Matsa and Miller (2013)'s evidence of a post-compliance decline in profitability, and by Bøhren and Staubo (2014)'s evidence that the quota may have caused a surge in corporate legal conversions to avoid the quota constraint.

The prior evidence and associated narrative is important not the least because it suggests that modest changes in board composition (the exchange of one or two male directors for females) can cause a decline in firm value—even in a setting where share ownership is concentrated. Moreover, the quota-based evidence draws statistical power from two noticeable aspects: First, as it is driven by gender politics only, the quasi-experiment is truly exogenous to firm performance, thus avoiding endogeneity issues when testing for effects of board changes on firm performance (Bhagat and Jefferis, 2005; Coles, Daniel, and Naveen, 2008; Adams, Hermalin, and Weisbach, 2010). Second, since the quota narrowly regulates gender balancing, it provides an unusually clear causal link to subsequent changes in board characteristics.[3]

It is is precisely the power of the experimental setting that has led us to address puzzling questions raised by the earlier narrative. First, is it reasonable that forcing shareholders to replace two male directors with females in a five-member board could destroy equity values upwards of 20% (AD, above quote)? In the corporate finance literature more broadly, value losses of this magnitude have been reserved for dramatic corporate events such as product recalls and financial distress (Jarrell and Peltzman, 1985; Hotchkiss, John, Mooradian, and Thorburn, 2008). Second, in Matsa and Miller (2013), the estimated reduction in return on assets (ROA) appears similarly outsized. Could it be that their coefficient estimate is influenced by firms' differential responses to the contemporaneous financial crisis rather than the quota *per se*? Third, AD document a significant

[2] *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).

[3] In contrast, the 2002 Sarbanes Oxley Act (SOX) in the US responded to performance scandals (such as the Enron scandal) and it mandated complex governance changes ranging from new internal control systems to enhanced director fiduciaries (Chhaochharia and Grinstein, 2007; Duchin, Matsusaka, and Ozbas, 2010).

reduction in board CEO experience. However, what is the change in board CEO experience if we eliminate the numerous single-person firms that are reflected in directors' curriculum vitae—the source of AD's measure of experience? Fourth, while a corporate legal conversions to avoid the quota requires the firm to delist from the OSE, is it reasonable to include, as AD do, delistings associated with complex transactions such as mergers and acquisitions (M&As) and going private transactions in the conversion count?

In order to provide satisfactory answers to these and other questions, we successfully replicate the prior studies before significantly expanding the analysis, leading us to substantially alter the prior scientific narrative. Our replication serves to remove disagreement over the data and econometric methodology employed by the prior studies. As a result, it also allows us to unambiguously point out and correct the econometric issues that we show are driving the most important earlier conclusions. We show that simple econometric (not data) corrections eliminate the economic and statistical significance of AD's valuation estimates. All of the main empirical conclusions of this paper are shown to hold not only for our expanded data set, but also for the sample periods of the earlier studies. In other words, our new empirical conclusions and associated narrative receive extraordinarily robust support in the data.

We begin by providing a comprehensive event study of the impact of major quota news events on stock returns of OSE-listed ASA. As recommended by a long series of papers,[4] we estimate abnormal stock returns using a portfolio approach, which adjusts the standard errors for the contemporaneous cross-correlation of stock return. These portfolio estimates produce statistically insignificant abnormal returns to the major quota-related news events, including the single event studied by AD (February 22, 2002). Cross-sectional regressions with firm announcement-returns as dependent variable and firm and board characteristics as regressors further corroborate this conclusion. When we similarly adjust AD's abnormal return estimate for cross-correlation, their abnormal returns also become statistically insignificant (whether using AD's sample of firms or our larger data-set). We end our analysis of stock return performance by forming multi-year long-

[4]E.g., Brown and Warner (1980, 1985), Schwert (1981), Kothari and Warner (2007) and Kolari and Pynnönen (2010)

short portfolios (long in firms with all-male boards and short in the remaining firms). None of these portfolios show statistically significant risk-adjusted performance.

Next, we implement AD's instrumental variable (IV) test for quota-induced valuation effects in Tobin's Q. Responding to the government's quota proposal in March of 2002, firms began changing their percentage of female directors already in 2002. In fact, in that year, as much as one third of the 146 listed ASA in 2001, including five large government-controlled ASA, subsequently reduced their shortfall of female directors. The final quota law came in 2005, requiring full compliance by the end of 2007. Over this six-year period, firms endogenously adopted to the (expected) quota constraint—some early and others late.

Citing the methodology in Stevenson (2010), AD's IV test specification is designed to eliminate such endogenous compliance activity in the first IV stage. However, in their implementation, they inexplicably use the distribution of female directors at year-end 2002 to identify pre-quota exogenous variation in female directors. We show that using the distribution at year-end 2001 (or earlier) alone drives the AD estimates from indicating persistent value losses of upwards of 20% for all-male boards to become small and statistically insignificant. Moreover, replacing the fraction of female directors with the *number* of female directors, which is not standardized by the (endogenous) board size, there is no evidence of a significant impact of the quota on Q regardless of the choice of base year in the first-stage regression.

We then turn to the analysis in Matsa and Miller (2013), who examine the effect of the quota on operating profitability (ROA) using a difference-in-difference (DID) regression. Using 2007-2009 as their post-quota estimation period, they identify a significant reduction in ROA for ASA relative to private limited companies (AS), which are not regulated by the quota. Their specification effectively identifies a decline in ROA after 2006, and our contribution here is simply to examine whether this decline is robust to increasing the sample period beyond 2009, reducing a potential impact of the financial crisis on the ROA estimate. After identifying much the same decline in ROA using their sample period, we report two new findings. First, the reduction in ROA does not survive our extension of the sample period beyond 2009. Second, a year-by-year analysis of

ROA reveals that treated firms experienced significantly lower ROA in 2008 only, and significantly higher ROA than the control firms in 2013. Overall, our longer sample period, which helps smooth any short-term cash-flow effects of the financial crisis, leads us to conclude that the quota had an insignificant effect on the ROA of ASA.[5]

Following our analysis of market value and cash-flow effects, we turn to changes in key board characteristics, such as board CEO experience, director seat concentration and busyness and, finally, legal (ASA to AS) conversions. Our most important finding in this section of the paper concerns the likely change in board CEO experience. AD report a significant decline in director CEO experience, which they argue corroborates their estimate of a significant reduction in equity values. However, we find that ASA succeeded in maintaining what we label board "large-firm" CEO experience. Large-firm CEO experience refers to CEO positions in all ASA and the 1% largest AS by revenue (Large AS) over the past three years. CEO experience in this subset of the largest firms is almost certainly more relevant for the firm value of listed firms than small-firm CEO experience. It is only when we expand our definition of CEO experience to include CEO positions in *any* AS, that our DID regressions indicate a significant decline in board CEO experience. Since AD uses director curriculum vitae to identify CEO experience, their measure in large part reflects such small firms, most of which have two employees or less. In sum, while small-firm CEO experience declined after quota implementation, large-firm CEO experience did not, further corroborating our finding of a value-neutral effect of the quota.

Next, much as reported in the extant literature, we show that the average board size in our sample remains largely unchanged at five shareholder-elected directors after the quota. However, we do detect a narrowing of the board size distribution from four to five—and from six to five—directors, reducing the need for replacing male directors with females (see Table I below). However, we also show that these board changes did not affect seat concentration, nor did board busyness increase (Fich and Shivdasani, 2006; Field, Lowry, and Mkrtchyan, 2013). Adding Bertrand, Black,

[5]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." We leave formal tests of this interesting hypothesis to future research.

Jensen, and Lleras-Muney (2019)'s finding that the new female directors are more qualified than their female predecessors, our overall conclusion is therefore that the pool of qualified potential female directors was deep—apparently allowing shareholders to rebalance the boards without loss of market value.

Finally, we expand on Bøhren and Staubo (2014)'s conclusion that ASA converted to AS to avoid the quota restriction. As firms do not disclose their reason for converting, inferences as to whether conversions are quota-driven are necessarily indirect. Bøhren and Staubo (2014) find that the conversion likelihood decreases in the fraction of female directors and conclude that conversions are quota-driven. We make two contributions to this conversion debate. First, 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), the conversion likelihood that they estimate becomes statistically independent of the quota constraint. Second, abstracting from M&As and going-private transactions, which are almost certainly not driven by quota considerations, not a single *listed* ASA converted to AS between 2001 and 2009. In fact, the number of OSE-listed firms increased every year as firms were complying with the quota.

The rest of the paper is organized as follows. Section I reports on our quota event study, followed by sections II and III on changes in Tobin's Q and operating profitability, respectively. We present our evidence on changes in board CEO experience in Section IV.A, on other board changes in Section IV.B, and on legals conversions in Section IV.C. Section V concludes the paper.

I Quota-induced valuation effects: event study

In this section, we estimate the average stock market reaction to major quota-related news events and investigate the determinants of individual firms' abnormal returns. Moveover, we empirically resolve differences between the conclusion from our event study and that of AD. We end the section with a long-run calendar-time estimation of the abnormal return to a long-short portfolio, where the long position is in the subsample of firms with the most binding regulatory constraint.

I.A The gender quota and its legislative timeline

Under Norway's codetermination law, shareholders elect one set of directors and employees another—up to one third of the board. Since the quota applies to shareholder-elected directors only, they are the exclusive focus of this paper. Directors are nominated by an independent committee and typically for a term of two years. Shareholder concentration is high. In the population of OSE-listed ASA, the largest shareholder owns on average 37% of the voting stock (median 31%). Thus, director elections are substantially influenced by shareholders.

Table I shows the mandated fraction of female directors and how it depends on board size, defined as the number of shareholder-elected directors. The required fraction of women 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 I 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 varies with board size, the firm can affect $Shortfall$ either by replacing male directors or by changing board size.

In 2001, no ASA board was gender-balanced and, as shown in Figure I, 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 retained all five males and appointed three new females to fill the quota. As discussed further in section IV.B below, 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 cost of expanding board size to eight directors places an upper bound on the expected shareholder-borne cost of the quota.

Turning to the time line, the gender quota legislation started by the quota receiving unexpected government support in 2002 and ended when all ASA complied at their 2008 annual shareholder meetings. We focus our event study on five news event dates listed in Table II, 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). Over the following week, the parliamentary members of the Minister's political party publicly reiterated the party's negative stance to the quota.

Given this opposition to a gender quota, it came as a further surprise when, on International Women's day (March 8, 2002), the coalition government unexpectedly proposed a gender quota legislation (event 2). Since the major opposition party also was in favor of a quota, this implied that the quota now had broad political support. Taking a lead, Cabinet promised compliance by government-owned firms within one year. The proposal was submitted to Parliament on June 16, 2003 (event 3). It contained a sunset provision, cancelling the quota if firms complied voluntarily by 2005. As expected, this law 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. To take effect, the quota had to be mandated by Cabinet once the sunset provision expired.

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 was resolved when Cabinet mandated the law on December 9 (event 5), without any special sanction for quota non-compliance. As a result, the sanction became forced liquidation—the ultimate penalty for violation of Norwegian corporate law.

I.B Portfolio estimation of abnormal returns

We explore the five events in Table II individually and jointly using daily stock returns of 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 narrow 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. All of our conclusions are robust to changing the portfolio weights to value-weights or, as recommended by Kolari and Pynnönen (2010), weights inversely proportional to the standard deviation of each individual sample firm's residual ε in the return generating process (Eq. (1) below).[6]

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:

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

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

[6] 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. The SUR approach produces identical coefficient estimates to our portfolio approach. However, it is less efficient because it substantially increases parameter estimation error (Sefcik and Thompson, 1986).

daily abnormal portfolio return over the two event days, and the two-day cumulative abnormal return is $CAR_k(-1,0) = 2AR_k$.[7]

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.[8].

Table III 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 III 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 I) 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 III

[7] 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).

[8] While not 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).

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 III 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.

I.C Determinants of announcement returns

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:

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

In addition to $Shortfall$, the vector of controls **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 IV.

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 V. 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 III and V fail to reject the null hypothesis of a zero valuation effect of the gender quota for OSE-listed ASA.

I.D Long-run abnormal returns

Using monthly returns from February 2002 (the first major quota event) through April 2008 (when all firms were in full compliance), we estimate the long-run abnormal return parameter α in the following return generating process:

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

Here, 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.

Table 3 shows α estimates for three alternative equal-weighted portfolios. In columns (1)-(3), the return generating process is Eq. (3) above, while columns (4)-(6) add a global momentum risk factor (MOM). The first portfolio, $Zero_{2001}$, contains an average of 98 OSE-listed ASA with zero female directors (all-male boards) in 2001. The second portfolio, Pos_{2001}, contains an average of 32 firms with at least one female director in 2001, while the third portfolio, $Zero$–Pos, is long in $Zero_{2001}$ and short in Pos_{2001}. The abnormal performance parameter α is statistically insignificant from zero for all three portfolios.

I.E Revisiting AD's event study

AD 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 VII 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. Since they report p-values only, the standard errors for the average $CAR^{AD}(-2,2)$ shown in parentheses are computed by us from their reported p-values.

As listed in Column (1) of Panel A, the sample average $CAR_i^{AD}(-2,2)$ is -2.57% with a reported $p = 0.001$ (implying a standard error of 0.757). Columns (2) and (3) show that the negative CAR^{AD} is concentrated in firms with zero female directors in 2001 ($Zero_{2001}$). In the last column, AD show that the difference between the sample of all-male boards and firms with at least one female director (Pos_{2001}) is -3.52% and highly significant using the p-test. AD also report sign and rank-sum tests for differences in median CAR^{AD} values. According to this test, the difference between median values in columns (2) and (3) is -1.6% with a marginally significant p-value of 0.054. In sum, AD concludes with a large negative market reaction to the February 22, 2002, quota event.

In Panel B of Table VII, 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 AD, 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.[9] 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 AD 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 VII, we generate correct standard errors by re-estimating AD's five-day industry-adjusted abnormal return using the portfolio time-series approach in Table III above. This portfolio estimation starts 252 days prior to February 22, 2002, has the equal-weighted

[9]While we use the matching procedure exactly as stated in footnote 9 of AD, the discrepancy may nevertheless reflect small differences in the selection of matching firms.

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. As AD, we also split the sample by the board gender composition in columns (2) and (3). Again, our average CAR estimates are similar to AD's, but statistically insignificant.[10]

Finally, in Appendix Table I, we examine AD's five-day event window day by day. Recall from Table II 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. Appendix Table I shows that this retraction—unreported by AD—has important implications for the interpretation of their five-day abnormal return estimate. While AD assume that their five-day abnormal return estimate reflects the pro-quota Friday announcement only, Panel A of Appendix Table I, which again uses AD's estimation methodology, shows that this assumption is wrong. In Panel A, the market reaction to the Minister's Friday announcement in support of the quota (day 0) is statistically insignificant even *without* correcting AD's standard errors for cross-sectional dependence.

[10]The following illustrates AD's understatement of the standard error of their abnormal return estimates. Recall from Table VII that 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 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

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

and so the test bias is

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

Using $5\sigma_{AR} = 3.250$ from Panel C and $\sigma_{CAR}\ independence = \sigma/\sqrt{94} = 0.780$ from Panel B implies an average ρ 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, AD's 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.

Moreover, Appendix Table I shows that AD'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, AD assign their negative abnormal return to the wrong event, leading to a misinterpretation of their own empirical result.

Switching back to our portfolio estimation methodology, shown in Panel B of Appendix Table I, 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 AD's five-day event window.

II The gender quota and Tobin's Q

In this section, we investigate whether the gender quota affects Tobin's Q. The main motivation is to provide additional perspective on our conclusion of a value-neutral effect of the gender quota. Since the Q analysis requires firms to be listed, we identify the effect of the quota-constraint using each firm's value of $Shortfall$. 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 Tobin's Q (henceforth Q) in subsequent periods.

However, even if the quota is exogenous, the timing of compliance may not be, as firms can differ in their ability to attract qualified female directors. As a result, the firms that comply early may have found it less costly to do so than late compliers. Since the quota news announcements in 2002 increase the likelihood of a quota law—but do not resolve who are likely to comply early or late—later compliance decisions may convey new information to the market and thus affect Q.

We begin with a reduced-form regression, followed by our implementation of the IV analysis designed by AD. As our results reverse theirs, we end the section by carefully reconciling the difference between the two studies.

II.A Reduced-form Q regression

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

$$Q_{it} = \alpha + \beta Shortfall_{it} + \theta_i + \tau_t + \epsilon_{it}, \tag{5}$$

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.

II.B Two-stage IV analysis of Q

We next employ the IV analysis designed by AD with a simple modification. Recall 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 announcements in February and March of that year. Specifically, since ASA general shareholder meetings (GM) typically take place in April/May, many firms started moving towards compliance with the expected future quota at the 2002 GM. In fact, 29% of the listed ASA included in the Q analysis, and with board information in years 2001 and 2002, reduced their shortfall of female directors in 2002. The reduction was the result of either increasing the number of female directors or reducing board size (we return to board changes in Section IV.B below). This early compliance activity implies that 2001 is the last year in which the board gender composition of ASA firms is exogenous to the quota.

The IV analysis is designed to eliminate effects of firms' endogenous compliance timing over the six-year compliance period, from 2002 to 2008. In the first stage, the firm's female director shortfall, $Shortfall_{it}$, is regressed on its exogenous shortfall in T_0 =2001, $Shortfall_{iT_0}$ interacted with year dummies, D_t, as follows:

$$Shortfall_{it} = \alpha + \beta_t D_t Shortfall_{iT_0} + \theta_i + D_t + u_{it}, \tag{6}$$

where θ is firm fixed effects. The predicted value of Eq. (6), $\widehat{Shortfall}_{it}$, extrapolates the firm's exogenous shortfall in 2001 with the market-wide trend, thus eliminating firm-level endogeneity in the appointment of female directors. In the second stage, Q is regressed on the predicted shortfall:

$$Q_{it} = \alpha + \beta \widehat{Shortfall}_{it} + \theta_i + \tau_t + \epsilon_{it}. \tag{7}$$

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

The first-stage regression estimates are shown in Panel B of Table VIII, 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.

Beginning in 2005, listed firms were required to report using the International Financial Reporting Standard (IFRS), switching from Norwegian General Accepted Accounting Principles (NGAP). This switch, which affected the balance sheets of all sample firms, raises a concern that the change in reported total assets is correlated with the pre-quota shortfall of female directors. However, while not tabulated, adding firm characteristics such *Total assets*, *Leverage* (ratio of book value of total debt to total assets), and *Board size* as control variables in the second stage of the IV test does not change our inferences. In sum, we conclude that Q is unrelated to the shortfall of female directors, also when instrumented with the market-wide time trend.

II.C Revisiting AD's analysis of changes in Q

While our coefficient estimate on $\widehat{Shortfall}$ is an insignificant 0.75 in column (1) of Table VIII, AD report a statistically significant coefficient estimate of -1.94 using a similar test design (Panel A of their Table IV). The main difference between the two test designs lies in the choice of exogenous base year T_0, where AD select T_0=2002. In fact, when we use T_0=2002 in our analysis, we produce a near-identical statistically significant coefficient estimate of 1.91 in Column (5) of Table VIII.

The sign of AD's coefficient estimate is negative because they base their analysis on the percent female directors, while we use the near-inverse variable $Shortfall$. Moreover, AD employ industry-adjusted Q, which we do not (Gormley and Matsa, 2014), and their sample is smaller than ours (603 firm-years versus 820). Notwithstanding these differences, it is the choice of T_0=2002 that drives the difference between the coefficient estimate of 0.75 in Column (1) of Table VIII and AD's 1.94. In fact, as shown in Column (7) of Table VIII, when we select T_0=2001 and use AD's estimation period, 2003-2009, our coefficient estimate is again small (0.69) and statistically insignificant.

Recall that firms began increasing female board participation in response to the quota already in 2002. Figure II illustrates the impact on the first-stage instrumentation of this early compliance activity. Each line in the figure is a firm-specific predicted compliance pathway $\widehat{Shortfall_i t}$ from

2003 through 2009 (AD's estimation period). In Panel A, the base year is $T_0 = 2001$ (this paper's specification), while in Panel B, $T_0 = 2002$ (AD's specification).[11]

Notice the increased cross-sectional dispersion of the pathways from Panel A to Panel B. This increase is caused by the early compliance activity in 2002. In particular, as some firms fully complied with the quota in 2002, Panel B of Figure II includes a pathway originating at zero shortfall in 2002. As discussed above, this pathway includes government-controlled ASA, which suggests a potential economic explanation for why the choice of base year has such dramatic impact on Q. To see why, note first that AD's negative impact on Q is primarily driven by the years 2007 and 2008 (see AD's Table IV, Panel B). Second, Beuselinck, Cao, Deloof, and Xia (2017) show that government-controlled listed companies perform relatively well during financial crises (perhaps due to an implicit government financing guarantee). This raises the possibility that AD's conclusion of a negative quota-induced impact on Q is driven *not* by the change in the percentage of female directors but by the relative impact of the financial crisis on government-controlled ASA.

Appendix Table II replicates Table VIII with the five government-controlled firms eliminated from the sample. These five firms, which are among the largest OSE-listed companies in terms of market value and number of employees, were under particular political pressure to comply early with the government's new quota policy and most did so already in 2002. As Appendix Table II shows, without the government-controlled firms in the sample, the coefficient on $\widehat{Shortfall}$ is statistically insignificant in all of the specifications. This is true also for Column (5), where the base year is 2002 as in AD. Thus, without the government-controlled firms in the sample, the conclusion is the same whether using year 2002 or an earlier year as the base-year for the IV test: a statistically insignificant impact of the quota law on Q.

Finally, columns (6) and (8) of Table VIII show that, regardless of the choice of base year T_0 (as 2002 or 2001), the second-stage regression coefficient for $Shortfall_{Number}$ is statistically insignificant. Thus, AD's main conclusion also hinges on using the *fraction* of female director

[11] The number of unique ASA in Figure II is 227, while the number of distinct pathways is much smaller: 13 in Panel A and 12 in Panel B). This is because firms with the same initial shortfall and board size have the same estimated pathway until 2008.

shortfall (not the female shortfall number) as the main regressor. We conclude from Table VIII and our robustness tests that the quota did not significantly impact the Q-values of OSE-listed firms.

III The gender quota and operating profitability

The analysis up to this point focuses directly on potential effects of Norway's forced gender balancing on firm value. In this section, we examine potential cash flow effects through ROA. While the *ex ante* expected abnormal stock return is zero, there is no similarly strong counterfactual prediction for changes in ROA. However, with a properly specified DID regression, evidence of a treatment effect on ROA is interesting as it may help indicate the existence of treatment effects in terms of firms' underlying production and investment policies.

Using the IV approach discussed in Section II.C above for changes in Tobin's Q, AD also test for changes in ROA and find no statistically significant effect. However, using a more powerful DID regression specification, Matsa and Miller (2013) report a post-quota (2007-2009) decline in ROA for treated firms. In this section, we primarily test if this negative effect on ROA survives extending the sample period beyond 2009. This test is of interest given the potential for a differential cash-flow effect of the financial crisis itself on the treated and control firms over Matsa and Miller (2013)'s short three-year post-quota window.

In our tests below, we use DID regressions with listed and unlisted ASA as treated firms and the 1% largest AS by revenue (henceforth "Large AS"—all unlisted) as control firms. As we show later (Section IV.C), we find that legal conversions from ASA to AS are statistically unrelated to the quota constraint, which helps validate the use of AS as control firms.[12] We begin with a description of the Large AS control firms before explaining our DID test results and how these relate to the findings reported by Matsa and Miller (2013).

[12]Conversions, which are executed through a bylaw change, occur regularly throughout our entire sample period, and firms rarely disclose their rationale for converting. Further discussion in Section IV.C below.

III.A Sample description of treated and control firms

Table IX reports bi-annual firm and board characteristics for ASA and Large AS over the period 1998-2013. Again, Large AS is the 1% largest AS by revenue. Using the *Brønnøysund Register Centre* data, for a firm-year observation to be included in the sample, the firm must have sufficient accounting information.[13] Moreover, we exclude subsidiaries not reporting consolidated accounts. These restrictions result in 1150 unique ASA (listed and unlisted) and 3532 unique Large AS.

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 and largely time-invariant 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.

Notice also that the quota appears to have had only a negligible spill-over effect on the gender representation of Large AS boards. Column (4) of Table IX 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.[14]

In the DID analyses, we require firms to have non-missing values for all board control variables and at least two observations during the sample period 2001–2013. Moreover, the DID regressions exclude financial firms because they were required to be ASA until 2007. Finally, as do MM, we recognize that the choice to be an ASA versus AS may be tied to the gender quota constraint itself (discussed further in Section IV.C below). We therefore exclude firms that convert their

[13] Specifically, we require the firm to have positive total assets and revenue, non-negative long-term assets, current assets, long-term debt, and short-term debt, as well as assets$\geq$cash and total assets$\geq$working capital (defined as current assets$-$current debt).

[14] This modest increase is similar to the general trend throughout western economies. For example, by 2013, the fraction of female directors was on average 18% in EU large publicly traded firms and 17% in US Fortune 1000 firms. Source: *http://ec.europa.eu./justice/gender-equality*, and *https://www.2020wob.com*.

legal status from ASA to AS and vice versa over the sample period (we relax this constraint in Appendix Table III, discussed below). With these additional restrictions, the final DID regression sample comprises 559 unique ASA and 1865 unique Large AS, for a total of 13,779 firm-years.

III.B ROA test results

We use the following OLS panel regression to test for quota-induced effects on ROA for ASA (treated firms) versus Large AS (control firms), with data from the period 2001-2013:

$$ROA_{it} = \gamma_0 + \gamma_1 ASA_i * Comply_t + \gamma_2 \mathbf{X_{i,t}} + \theta_i + \tau_t + \epsilon_{it}, \tag{8}$$

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.

The vector $\mathbf{X}_{it}$ contains the following firm characteristics (defined above and in Table IV): *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*. The inclusion of *Size* and *Leverage* helps to control for the concern that the switch by listed firms to IFRS in 2005 may drive the results. *Board CEO experience* is the fraction of directors with *Large-firm CEO experience*, which is a binary variable taking the 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 because 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 X. The first four columns use the full sample period (2001-2013). In Columns (1) and (2), 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, using the entire sample period, there is no evidence that the forced gender balancing affects operating profitability. Moreover, columns (3) and (4), which decompose *Comply* into year-by-year effects ($ASA * 2008$, $ASA * 2009$, etc.), show that there is no negative effect on ROA in any of the years and a significantly *positive* effect in 2013.

The conclusion from columns (1)–(4) of Table X contrasts with the negative treatment effect identified by Matsa and Miller (2013). In columns (5)–(10) of Table X, we use our sample to test whether the different estimation periods in the two papers is what drives our different estimate of the treatment effect.[15] In Columns (5) and (6), we use Matsa and Miller (2013)'s shorter sample period, 2003–2009. This produces a coefficient estimate for $ASA*Comply$ that is negative (–0.025 and –0.023 in columns (5) and (6), respectively) and significant at the 5% level. These coefficient estimates are close to the –0.027 reported by Matsa and Miller (2013) when using a matched sample of Norwegian AS as control firms (columns 1–2 of their Table 3).

In the last four columns of Table X, we eliminate the first two pre-quota years from the sample period. Columns (7)–(8) reveal that the insignificant ROA effect in columns (1)–(2) is robust to eliminating years 2001 and 2002 from the estimation. Moreover, decomposing *Comply* into year-by-year effects in columns 9–10 show that ROA is significantly lower for treated firms in 2008 only. Thus, it appears that the negative effect of gender balancing found by Matsa and Miller (2013) on operating profitability is dependent on the pre-quota sample period starting in 2003, is short-term, and is unique to the financial crises year of 2008.

Recall that Table X excludes from the sample any firm that switches legal form over the sample period (ASA to AS or vice versa). Since separate analysis (Section IV.C below) shows that the propensity to switch between ASA and AS is independent of the board gender composition, this sample exclusion should not matter for the overall conclusion. This is indeed confirmed in

[15]Matsa and Miller (2013) study differs from ours in terms of sample period (2003-2009 versus our 2001-2013), but also in terms of the number and selection of treated firms (104 listed ASA in their main analysis versus our 559 listed and unlisted ASA) and selection of control firms (matched AS versus our Large AS). Matsa and Miller (2013) exclude firms in the petroleum industry and use a different data source (Orbis), which cover fewer Norwegian firms than our population data. They also use listed firms in other Nordic countries as controls, which we do not.

Appendix Table III, where we rerun the regressions in Table X with switchers included (classifying their legal status as that of their first year in the sample). As shown, the sample expansion drives the coefficient for $ASA * Comply$ to become insignificant also for the 2003-2009 sample period (columns 5–6), thus reinforcing our conclusion of a statistically insignificant ROA treatment effect.

IV On the supply of qualified female directors

Contrary to the extant literature, our empirical evidence above—on stock the market reaction to quota news announcements and time series changes in Tobin's Q and ROA—all point to a neutral effect of the gender quota on equity values and firm profitability. In this section, we investigate whether post-quota changes in specific board characteristics corroborate this central finding. Essentially, we are asking whether the supply of qualified female directors appear to have been sufficiently high for the quota to have been a low-cost event for shareholders. To this end, we present new descriptive evidence ranging from quota-induced changes in board CEO experience and director seat concentration and busyness, as well as on the propensity for ASA-to-AS legal conversions. For reasons that we carefully explain, the evidence in this section continues to challenge extant research on the effects of Norway's gender quota.

IV.A Did board CEO experience decline?

Director CEO experience is generally viewed as central for board effectiveness and thus valued by investors (Fich, 2005; Fahlenbrach, Low, and Stulz, 2010; Kang, Kim, and Lu, 2018). As we show in Column (6) of Table IX above, the fraction of female CEOs in ASA and Large AS is low: only 5% or less.[16] It is worth pointing out that this low percentage of female CEOs does not necessarily imply that the overall board CEO experience will decline when complying with the gender quota. The reason is that boards may chose to replace male directors without CEO

[16] Bertrand, Black, Jensen, and Lleras-Muney (2019) report a small increase in female ASA CEOs. Column (6) shows that this increase is largely driven by unlisted ASA: the fraction of female CEOs increases from 2% in 1998 to 5% for listed ASA and 10% for unlisted ASA in 2013. Column (5) of Table IX also shows an increase in female chairs between 1998 and 2013, up from 2% to 9% for listed ASA and from 1% to 15% for unlisted ASA.

experience with females or expand board size to make room for female directors while retaining experienced males. Thus, the question of whether the quota caused board CEO experience to decline is an open empirical issue.

The following analysis is novel in that it constructs two measures of board CEO experience. The first is *Large-firm CEO experience* which, as defined above, is restricted to ASA and Large AS. The second is *Small-firm CEO experience*, which records CEO experience in ASA and *all* AS, i.e., including the remaining 99% of the roughly 100,000 AS in an average year. We explicitly distinguish CEO experience in large and small firms because the population of AS is overwhelmingly dominated by tiny firms: 46% of all AS have at most one employee, 58% have at most two, and 90% have at most ten. The annual number of employees averages 657 for listed ASA, 209 for unlisted ASA, and 45 for Large AS, all of which are included in our measure *Large-firm CEO experience*.[17] In sum, an individual director's *Large-firm CEO experience* is undoubtedly of greater value to the board of an OSE-listed ASA than is *Small-firm CEO experience*.

As shown in Column (7) of Table IX, board large-firm CEO experience, i.e., the fraction of director's with *Large-firm CEO experience*, averages 17% in listed ASA, 15% in unlisted ASA, and 14% in Large AS.[18] In contrast, Column (8) shows that board small-firm CEO experience is substantially higher, averaging 53% in listed ASA, 47% in unlisted ASA, and 44% in Large AS.

Figure III plots the time series of board large-firm CEO experience in Column (7) pooled across listed and unlisted ASA, as well as that of male and female directors. The figure confirms a higher level of CEO experience among male directors. More important, there is an increase in male director CEO experience around quota compliance, suggesting that boards made an effort to retain male director CEO experience. This retention attenuates the tendency in Figure III for board large-firm CEO experience to decline.

In columns (1) and (2) of Table XI, we use Eq. (8) to test whether the quota caused changes in large-firm CEO experience for ASA boards relative to the control group of Large AS. The

[17] The frequency distribution of ASA is skewed towards firms with at most 50 employees.

[18] Fahlenbrach, Low, and Stulz (2010) report that 10% of outside directors in listed US firms are current outside CEOs.

vector $\mathbf{X_{it}}$ of controls contains *Firm age*, *Size*, *Leverage*, and *Largest owner* (replacing *Size* with *Total Assets* does not change our inferences). The coefficient estimates of γ_1 on $ASA * Comply$ are statistically insignificant, whether or not firm controls are included. This fails to support the hypothesis that ASA-board large-firm CEO experience falls relative to that of Large AS after quota compliance.

Turning to our second measure of board CEO experience, *Small-firm CEO experience*, the DID regressions in columns (3) and (4) of Table XI shows a significantly negative coefficient estimate for $ASA * Comply$. Since board large-firm CEO experience does not decline (columns 1 and 2), this decline is overwhelmingly dominated by experience in tiny AS. This finding helps clarify the decline in CEO experience documented by AD. Relying on director bios in company annual reports, which record all types of CEO experience, AD's measure is more closely related to *Small-firm CEO experience* than to *Large-firm CEO experience*.[19]

IV.B Other board changes

The previous section shows that boards, in complying with the quota, successfully maintained large-firm CEO experience. In this section, we examine whether this success necessitated changing other board characteristics, such as size, seat concentration, and director busyness. If qualified females are in short supply, an efficient response to the quota constraint may require a change in these characteristics.

Using board size as dependent variable, columns (5) and (6) of Table XI test whether the quota caused the average ASA board size to change differently from that of Large AS. 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 weak evidence of a decline in ASA board size.

To better understand this decline, Panel A of Figure IV plots the board-size frequency distri-

[19] AD's definition is broader than our *Small-firm CEO experience* as it includes "work experience as CEO or owner" and does not restrict the look-back period. In their sample, 66% of male directors and 40% of female directors have CEO experience (online Appendix Table II), which is close to the small-firm CEO experience in column (8) of our Table IX.

bution for 555 ASA in 2001 and 395 ASA in 2008. Interestingly, while Figure I demonstrates that the average board size remains constant at five directors, Panel A of Figure IV shows a narrowing of the distribution, reflecting a shift from four to five board members and from six to five—with a somewhat greater downward shift. This narrowing helps illustrate why the DID regression coefficient estimate of $ASA * Comply$ comes out negative. Moreover, the narrowing may be motivated by a desire to minimize the quota constraint: Increasing the board from four to five members allows a firm to appoint two females while retaining three (rather than 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.

Next, we turn to changes in board seat concentration and director busyness. We use a definition of busy directors—holding three or more board seats (here in ASA and Large AS)—found in the extant literature (Fich and Shivdasani, 2006; Field, Lowry, and Mkrtchyan, 2013). Panel B of Figure IV plots the frequency distribution of the number of board seats in ASA and Large AS held by male and female ASA directors in 2001 and 2008. ASA directorships are highly dispersed—almost three quarters of individual directors hold only one board seat—and the distribution is largely similar for male and female directors. The 2001 and 2008 proportions of busy directors are 12% and 13%, respectively, for male directors and 9% and 11%, respectively, for female directors. Thus, the level of female director busyness largely increased toward that of male directors.

In columns (7) and (8) of Table XI, we use our DID regression setup to test whether there is a quota-induced change in board busyness, defined as the fraction of busy directors. The coefficient estimate for the interaction variable $ASA * Comply$ is negative and significant in both columns. That is, overall ASA board busyness decreases post-quota relative to unregulated Large AS.

IV.C Legal conversions from ASA to AS

In this section, we follow the lead of Nygaard (2011) and Bøhren and Staubo (2014) and examine whether legal conversions between ASA and AS reflect the quota-constraint. Conversions, which are executed through a bylaw change, occur regularly throughout the sample period. While firms rarely disclose the rational behind the conversion decision, only ASA can list on OSE. Thus, a straightforward hypothesis for conversion from ASA to AS is an aborted plan to raise public equity. For such firms, there are few, if any, other benefits from remaining an ASA. However, there are certain costs, including having to register the shares with Norway's securities registry (VPS), as well as some additional corporate governance restrictions.[20]

The quota may have pushed some firms to convert from ASA to AS in order to avoid gender balancing. This hypothesis predicts that conversion likelihood is increasing in the female director shortfall. To test this proposition, we collect information on conversions from *Brønnøysund Register Centre*. Since conversion can be effectuated by a simple shareholder resolution, we rule out much more complex and costly transactions, such as mergers as bankruptcies as being driven by quota avoidance. Obviously, if an ASA is acquired and left as a free-standing AS entity, the legal conversion is a by-product of and not the underlying motivation for the transaction. Domestic M&A and bankruptcy filings are identified in the *Register*, and we search news and press releases to identify acquisition by foreign buyers. Finally, we exclude financial ASA because they were required to be incorporated as ASA until 2007.

Column (1) of Table XII 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 increases each year from 140 in 2004 to 195 in 2008. Column (2) lists the total number of firms that exit the ASA legal form each year. Column (3) shows annual exits due to M&A (takeovers by strategic and financial buyers and going-private transactions) and bankruptcy. These complex transactions are almost certainly not driven by quota avoidance and are therefore excluded from the count. As shown in Column (4), there are no other conversions by listed ASA. Over the entire compliance period, whatever the perceived cost of the gender quota, it was evidently lower than the benefits

[20] OSE-listed firms must report according to 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. Moreover, unlisted ASA are required to report trades in its shares to the Norwegian Securities Register (VPS) and insiders must report their trades to the board. The fraction of non-voting shares is limited to 50% for ASA, with no restriction for AS.

of remaining publicly listed.[21]

In Panel B of Table XII, Column (4) shows a steady number of conversions by unlisted ASA for reasons other than M&A and bankruptcy. The question is whether these other conversions, which spike in 2006–2007, are driven by the quota constraint. To formally test this hypothesis, we estimate the following logit model for the conversion likelihood with the female director shortfall as explanatory variable:

$$Convert_{it} = \alpha + \gamma_1 Shortfall_{it} + \gamma_2 \mathbf{X}_{it} + \kappa_i + \tau_t + \epsilon_{it}, \tag{9}$$

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 the logit estimation are in Columns (1)-(4) of Table XIII. 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 XIII are new to the quota debate, Bøhren and Staubo (2014) estimate a different conversion likelihood and conclude 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 (backfilling) for firms that convert. However, because firms drop out of the sample after conversion, the combi-

[21] *The Economist* got this wrong in an article headlined "Companies fled the [Norwegian] stock market as quotas were faced in" (November 15, 2014).

nation of backfilling the dependent variable and the time trend in female director representation mechanically produces 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. (9), 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.[22]

V Conclusion

Norway's pioneering board gender quota provides a unique quasi-experiment for identifying market valuation effects of exogenous changes in board characteristics. We address an important narrative emerging from prior research on this quasi-experiment, namely that the quota law caused an economically large decline in the market value of listed firms—possibly explained by an apparent post-quota decline in board CEO experience. Upon closer scrutiny, neither the earlier estimates of a negative quota-induced valuation effect nor the purported explanation for this negative impact survive our closer inspection. It is worth stressing that we show explicitly that our statistical inferences hold not only within our larger sample, but also after replicating several previous studies on their own terms before implementing the required econometric adjustments.

Table XIV conveniently summarizes the considerable number of new empirical results of this paper and how they relate to the most relevant prior studies. First, simple but necessary econometric adjustments produce statistically insignificant valuation effects whether we estimate abnormal stock returns to major quota-related news events (Panel A), long-short portfolio performance estimation (short in firms with pre-quota all-male boards), or IV tests for changes in Tobin's Q (Panel B). Second, post-quota changes in return on assets are statistically indistinguishable from zero, except in the financial crisis year of 2008 when there is a relative drop in the ROA of ASA (Panel C).

Third, with access to the population of directorships in ASA and AS, we show that board

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

CEO experience in ASA and the 1% largest AS does *not* decrease post-quota (Panel D). Earlier evidence of a decline in board CEO experience is instead driven by tiny firms (0-2 employees), which we argue hardly qualifies as relevant board experience for publicly traded companies. Board size, director busyness, and board seat concentration are also largely unchanged, which further supports the notion of a (perhaps surprisingly) deep supply of qualified female directors in the female population. Finally, we also show that, while legal conversions from ASA to AS occur routinely over the sample period, the impact of board gender composition on the conversion decision is statistically insignificant (Panel E).

The new empirical narrative emerging from our empirical results is that investors and firms alike viewed forced gender-balancing as a low-cost constraint on the free election of boards. This narrative may also help explain why several other western European countries have since decided to follow suit and adopted their own versions of board gender quotas—most recently in the state of California. The importance of our study is to show that, for a country like Norway, a board gender quota constitutes a low-cost regulation that is ultimately focused on promoting gender-neutral policies lower down in the corporate organization.

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Figure I
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*.

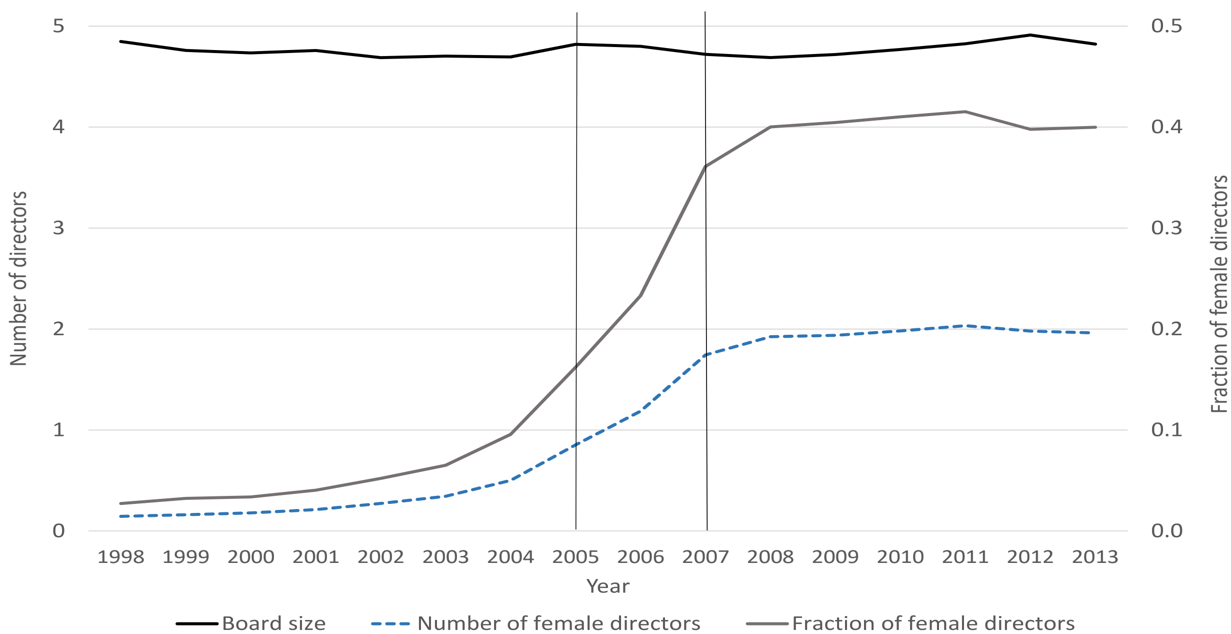


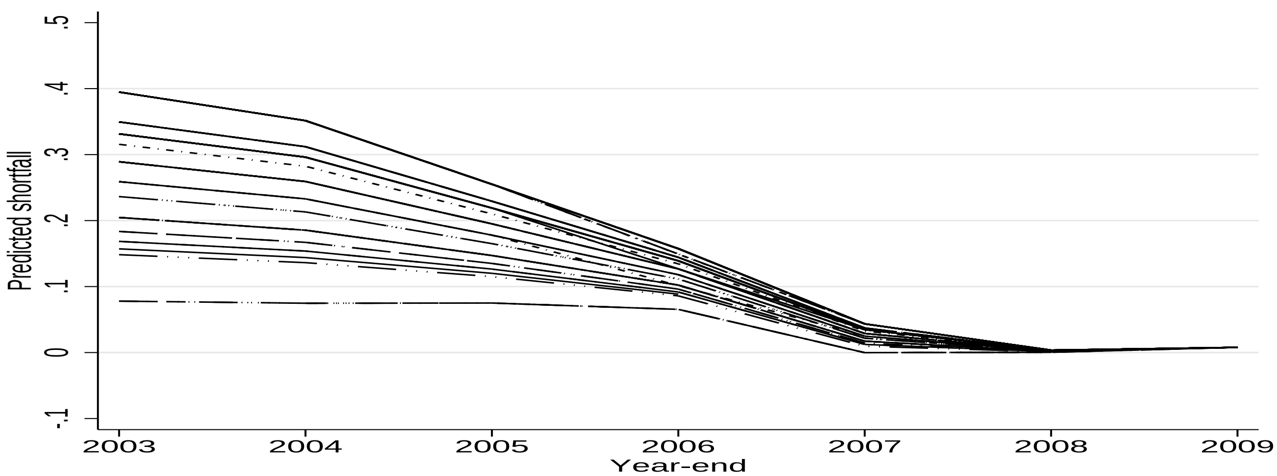
Figure II
First-stage instrumentation of female director shortfall

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, respectively, year and firm fixed effects. $Shortfall_{iT_0}$ 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 $T_0 = 2002$

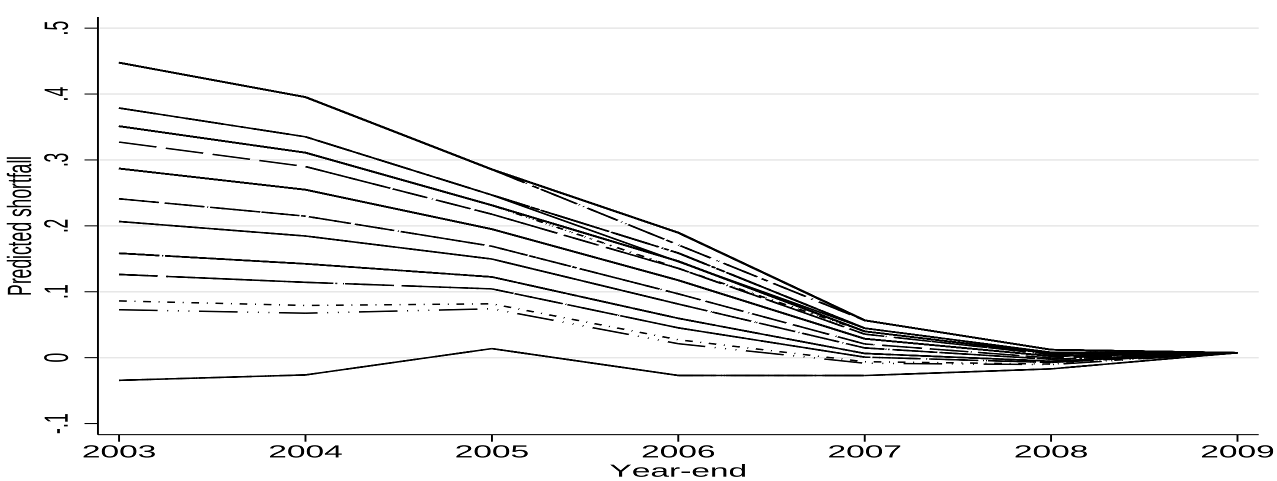


Figure III
Large-firm CEO experience of ASA directors and boards, 2001-2013

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*.

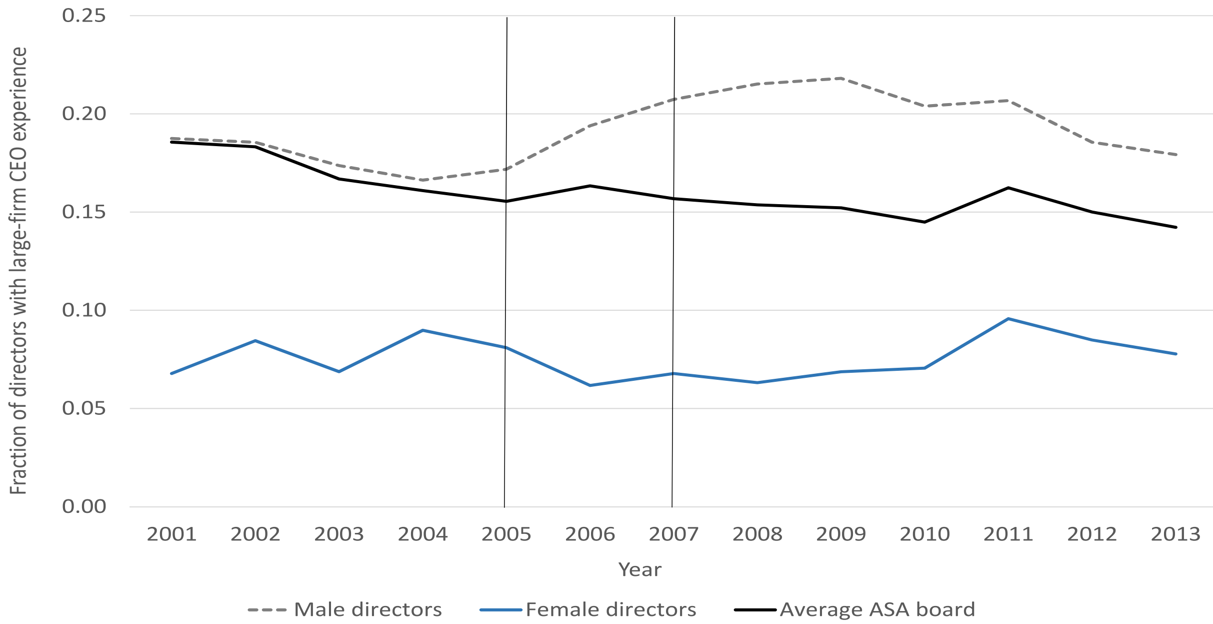
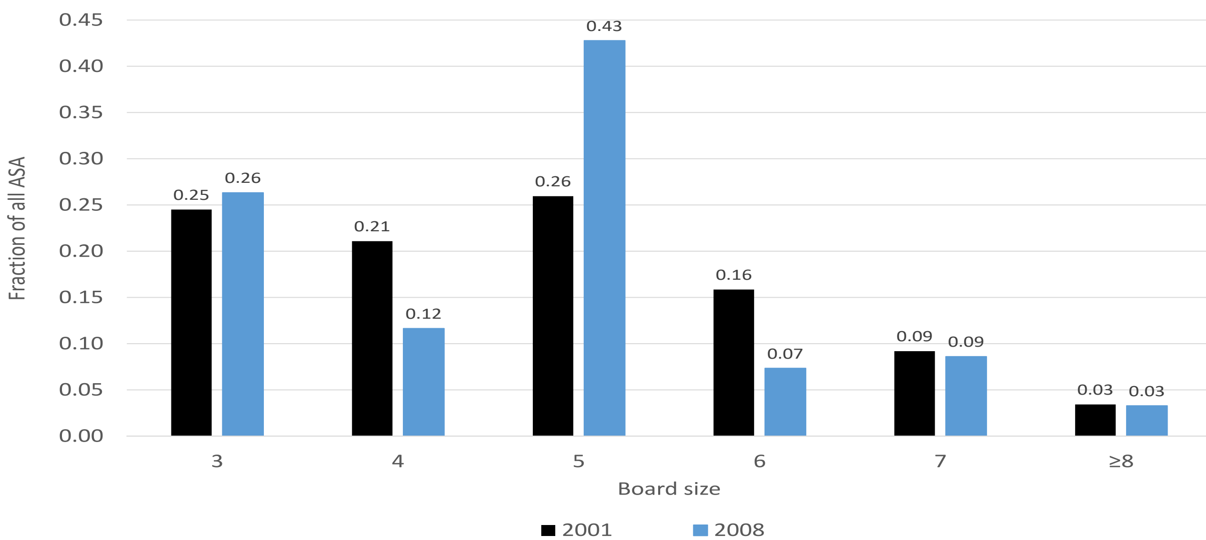


Figure IV
Frequency distribution of ASA board size and director board seats, 2001 and 2008

Panel A shows the frequency distribution of board size. Panel B plots the frequency distribution of the total number of board seats in ASA and Large AS (the top 1% AS by revenue) held by male and female directors. Five and more board seats are reported under 5+. The sample is 555 ASA (1938 male and 104 female directors) in 2001 and 395 ASA (909 male and 581 female directors) in 2008. Board data are from *Brønnøysund Register Centre*.

A: ASA board size



Panel B: Male and female ASA directors' board seats in ASA and Large AS

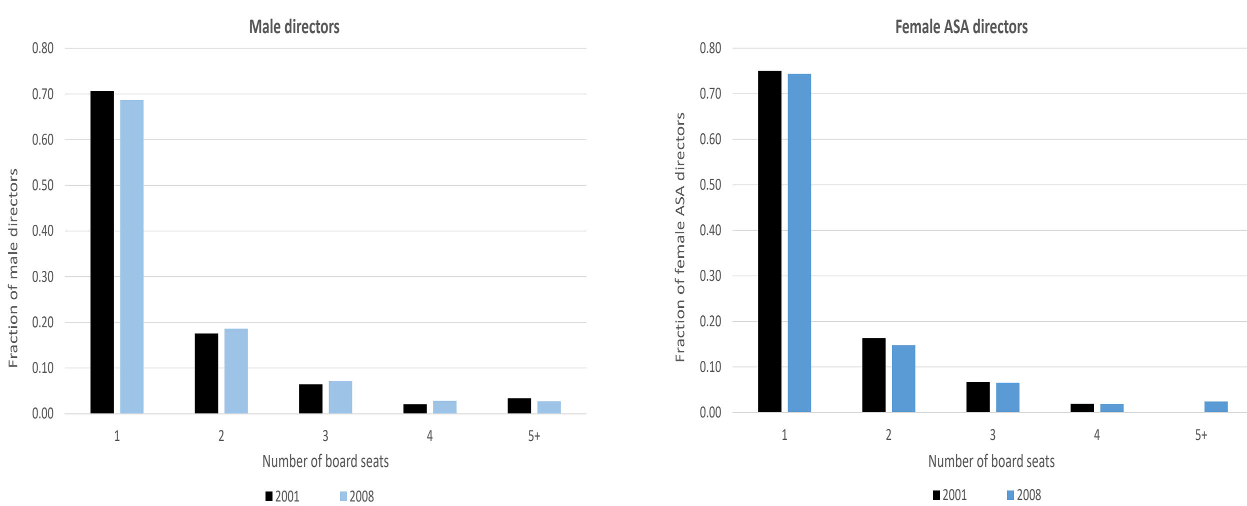


Table I
Female directors required by Norway's board gender quota

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	Required fraction of female directors	*Shortfall* on board with		
			1 female	2 females	3 females
	(1)	(2)	(3)	(4)	(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 II
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 III
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 II. 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 IV
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, and 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 the past three years.
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 I.
$Shortfall_{Number}$	Difference between the number of female directors required by the quota and that of the current board. See Table I.
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 V
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 II). For each firm i, the daily average abnormal return AR_{ik} is estimated for each event k using the regression model in Table III. 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 IV. 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 VI
Calendar time portfolio returns for zero and positive female representation in 2001

The table reports monthly abnormal stock returns for portfolios of listed ASA with zero or positive female representation in 2001, over the period February 2002 (start of quota legislative process) to April 2008 (full quota compliance). A $Zero_{2001}$ firm has zero female directors in 2001, while a Pos_{2001} firm has at least one female director in 2001. The monthly average number of firms in the $Zero_{2001}$ and Pos_{2001} portfolios is 98 and 32, respectively. In columns (1) - (3), the abnormal stock return is estimated using the following three factor return-generating process:

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

where r_t^e is the monthly USD-denominated stock return to domestic OSE-listed ASA, converted to USD using the monthly exchange rate, in excess of the current month's 3-month US Treasury bill. W^e is the monthly return on MSCI world stock market index in excess of the current month's 3-month US Treasury bill. SMB (size) and HML (value) are global risk factors from Ken French's web site. Columns (4)–(6) include an additional global momentum risk factor (MOM), also from Ken French's web site. Standard errors in parenthesis and significance levels are indicted by *** 1%, ** 5%, * 10%.

	$Zero_{2001}$ Portfolio (1)	Pos_{2001} Portfolio (2)	*Zero-Pos* Portfolio (3)	$Zero_{2001}$ Portfolio (4)	Pos_{2001} Portfolio (5)	*Zero-Pos* Portfolio (6)
α	-0.005	0.000	-0.005	-0.004	0.002	-0.005
	(0.007)	(0.004)	(0.004)	(0.007)	(0.004)	(0.004)
W^e	1.356***	1.354***	0.002	1.311***	1.277***	0.034
	(0.169)	(0.110)	(0.109)	(0.180)	(0.114)	(0.115)
HML	-0.368	0.074	-0.442	-0.313	0.166	-0.480
	(0.469)	(0.305)	(0.301)	(0.476)	(0.302)	(0.306)
SMB	0.927***	0.531**	0.396*	1.002***	0.659***	0.344
	(0.341)	(0.222)	(0.219)	(0.356)	(0.226)	(0.228)
MOM				-0.157	-0.265**	0.108
				(0.207)	(0.131)	(0.133)
R^2	0.533	0.705	0.058	0.537	0.722	0.067
Observations (months)	75	75	75	75	75	75

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

Panel A lists the average cumulative abnormal returns $CAR^{AD}(-2,2)$ reported by 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 II) 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 III, 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 VIII
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 I) 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 IV. 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 IX
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 IV 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 X
Quota-induced changes in operating profitability

The table reports coefficient estimates from the following 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 (all suppressed). The variables are defined in Table IV. 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)–(4), 2003–2009 in columns (5)–(6), and 2003–2013 in columns (7–10). Standard errors clustered by firm are reported in parenthesis. Stars indicate significance levels: *** 1%, ** 5%, and * 10%.

	2001-2013				2003-2009		2003-2013			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
$ASA * Comply$	0.004	0.012			–0.025**	–0.023**	–0.010	–0.001		
	(0.011)	(0.011)			(0.012)	(0.011)	(0.010)	(0.009)		
$ASA * 2008$			–0.026*	–0.022					–0.043***	–0.038***
			(0.014)	(0.014)					(0.014)	(0.014)
$ASA * 2009$			–0.001	0.001					–0.016	–0.013
			(0.015)	(0.014)					(0.014)	(0.013)
$ASA * 2010$			0.017	0.029*					0.004	0.017
			(0.016)	(0.015)					(0.015)	(0.014)
$ASA * 2011$			0.006	0.018					–0.007	0.007
			(0.020)	(0.018)					(0.020)	(0.017)
$ASA * 2012$			0.016	0.030*					0.004	0.020
			(0.019)	(0.018)					(0.019)	(0.017)
$ASA * 2013$			0.041**	0.050***					0.028*	0.040**
			(0.016)	(0.016)					(0.016)	(0.016)
Firm characteristics	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.024	0.109	0.026	0.111	0.025	0.136	0.018	0.123	0.022	0.128
N (firm-years)	13,779	13,779	13,779	13,779	7365	7365	11,531	11,531	11,531	11,531

Table XI
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 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 IV. 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 XII
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)	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 XIII
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 IV. 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 XII). 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

Table XIV

Summary of the effects of gender balancing on firm value, performance, CEO experience, and legal conversion

Each panel summarizes the samples and conclusions of the two leading empirical studies. Replication implies that this paper successfully generates the main conclusion of the other leading study. N1 is the number of unique treated firms, N2 is the number of firm-years, and sample period is the one used in the paper's main regression. CAR is cumulative abnormal stock return, IV is instrumental variable, AD is Ahern and Dittmar (2012), MM is Matsa and Miller (2013), and BS is Bøhren and Staubo (2014).

Study	N1, N2, (Sample period)	Main conclusion	Is earlier conclusion replicated here?	How does this paper change the earlier conclusion?
A. Firm value: CAR centered on February 22, 2002				
This paper	N1=143	CAR is statistically insignificant.	Yes, using either AD's sample with our data or AD's CAR estimation methodology.	Correcting AD's standard errors for return cross-correlation eliminates the statistical significance of AD's CAR
AD	N1=94	CAR is significantly negative.		
B. Firm value: Tobin's Q				
This paper	N1=239, N2=815 (2002-2008)	No significant effect of quota on Q.	Yes, using AD's methodology.	Correcting the 1^{st} stage of AD's IV test for endogeneity, by switching the base-year from 2002 to 2001, eliminates the statistical significance.
AD	N1=122, N2=605 (2003-2009)	Significantly negative effect of quota on Q.		
C. Firm performance: profitability (ROA)				
This paper	N1=559, N2=13,779 (2001-2013)	No significant effect of quota on ROA.	Yes, using MM's sample period.	The negative ROA-effect exists in the crisis year 2008 only and disappears when extending the post-quota period beyond 2009.
MM	N1=104, N2=3,116 (2003-2009)	Significantly negative ROA effect.		
D. Board CEO experience				
This paper	N1=559, N2=13,779 (2001-2013)	No significant change in CEO experience following quota.	Yes, using a diff-in-diff regression.	Small-firm CEO experience falls while Large-firm CEO experience (in ASA and 1% largest AS) does not. AD's results are driven by small firms (0-2 employees).
AD	N1=122, N2=605 (2003-2009)	Significant reduction in experience for the most affected firms.		
E. Legal conversions from ASA to AS				
This paper	N1=264 N2=880[a] (2001-2008)	Conversion decision is independent of board gender composition.	Yes, using BS' back-filled dependent variable.	Adding year fixed effects eliminates the backfilling bias in BS and renders board gender composition insignificant for the conversion decision.
BS	N2=1,560[b] (2000-2009)	Conversion likelihood increases with shortfall female directors.		

[a] N2 excludes listed ASA because of our evidence (Table XII) that no listed ASA converted to AS over the sample period for reasons other than M&A and bankruptcy.
[b] N2 includes listed ASA. N1 is unavailable from the original study.

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

The table repeats the estimation reported in Table VII for each day in the five-day event window (-2,2) around February 22, 2002 (event 1 in Table II). Panel A reports average daily AR for the sample firms in 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 somewhat 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]

Appendix Table II
IV regressions for Tobin's Q and female director shortfall, ex government firms

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 I) 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 IV. The sample is 234 OSE-listed ASA, 2002-2008 (columns 1-4) and 222 OSE-listed ASA, 2003-2009 (columns 5-8), excluding five firms controlled by the government. 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								
$\widehat{Shortfall}$	0.946		0.469		1.680		–0.145	
	(1.171)		(1.760)		(1.053)		(2.022)	
$\widehat{Shortfall}_{Number}$		–0.061		0.010		0.060		–0.367
		(0.189)		(0.227)		(0.152)		(0.339)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
F-statistic	17.28	17.05	18.75	20.18	15.66	15.54	16.02	14.66
N (firm-years)	780	780	691	691	785	785	755	755
B: First-stage IV regression for $Shortfall$								
$Shortfall_{T_0} \times D_{2002}$	0.560***	0.540***	0.537***	0.520***				
	(0.098)	(0.094)	(0.144)	(0.092)				
$Shortfall_{T_0} \times D_{2003}$	0.462***	0.424***	0.455**	0.414***	1.049***	0.963***	0.488***	0.415***
	(0.152)	(0.131)	(0.191)	(0.135)	(0.120)	(0.107)	(0.166)	(0.144)
$Shortfall_{T_0} \times D_{2004}$	0.442***	0.345***	0.537***	0.361***	0.943***	0.747***	0.463***	0.335***
	(0.128)	(0.110)	(0.153)	(0.121)	(0.142)	(0.117)	(0.147)	(0.126)
$Shortfall_{T_0} \times D_{2005}$	0.260**	0.182*	0.334**	0.251**	0.571***	0.471***	0.275*	0.167
	(0.117)	(0.107)	(0.148)	(0.118)	(0.146)	(0.119)	(0.140)	(0.125)
$Shortfall_{T_0} \times D_{2006}$	0.099	0.086	0.181	0.073	0.488***	0.349***	0.106	0.063
	(0.100)	(0.095)	(0.130)	(0.108)	(0.129)	(0.099)	(0.120)	(0.111)
$Shortfall_{T_0} \times D_{2007}$	0.098**	0.077*	0.055	0.017	0.258***	0.183***	0.100	0.054
	(0.049)	(0.045)	(0.070)	(0.044)	(0.079)	(0.056)	(0.064)	(0.056)
$Shortfall_{T_0} \times D_{2008}$					0.092	0.039*	0.003	–0.015
					(0.063)	(0.022)	(0.059)	(0.033)
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
F-statistic	81.15	81.70	82.94	72.63	87.05	63.81	47.53	45.53
N (firm-years)	797	797	705	705	794	794	764	764

Appendix Table III
Quota-induced changes in operating profitability

The table reports coefficient estimates from the following 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 (all suppressed). The variables are defined in Table IV. The sample comprises 559 ASA (treated firms) and 1865 Large AS (control firms), 2001-2013. Treated and control firms are identified based on their legal form in 2001 (columns 1-4) and 2003 (columns 3-10), or the first year they appear in the sample. We exclude financial firms, 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)–(4), 2003–2009 in columns (5)–(6), and 2003–2013 in columns (7–10). Standard errors clustered by firm are reported in parenthesis. Stars indicate significance levels: *** 1%, ** 5%, and * 10%.

	2001-2013				2003-2009		2003-2013			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
$ASA * Comply$	0.007	0.014			–0.018	–0.017	–0.004	0.003		
	(0.010)	(0.010)			(0.011)	(0.010)	(0.009)	(0.009)		
$ASA * 2008$			–0.023	–0.019					–0.037***	–0.033**
			(0.015)	(0.014)					(0.014)	(0.014)
$ASA * 2009$			0.006	0.005					–0.006	–0.006
			(0.014)	(0.013)					(0.013)	(0.013)
$ASA * 2010$			0.019	0.029**					0.007	0.019
			(0.015)	(0.014)					(0.014)	(0.014)
$ASA * 2011$			0.008	0.018					–0.003	0.010
			(0.019)	(0.017)					(0.018)	(0.016)
$ASA * 2012$			0.018	0.030*					0.008	0.022
			(0.018)	(0.016)					(0.017)	(0.015)
$ASA * 2013$			0.041***	0.048***					0.029**	0.039***
			(0.015)	(0.016)					(0.015)	(0.015)
Firm characteristics	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.024	0.108	0.026	0.111	0.024	0.136	0.018	0.124	0.021	0.128
N (firm-years)	13,875	13,875	13,875	13,875	7418	7418	11,634	11,634	11,634	11,634

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