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Abstract

We critically revisit and expand previous studies on the likely valuation effect of Norway's pioneering board gender-quota law. Most important, Ahern and Dittmar (2012) report a significantly negative average abnormal stock return, which they conclude is evidence of a large shareholder-borne cost of the quota constraint. We first show that they allocate their negative market reaction to the wrong event, which reverses their inference. We then show that their abnormal return estimate becomes statistically insignificant once we make the necessary adjustment for contemporaneous cross-correlation of stock returns. Furthermore, we provide new evidence on long-run abnormal stock returns, changes in Tobin's Q and operating profitability, and legal conversions, all of which corroborate that Norway's quota law most likely caused a statistically insignificant valuation effect on regulated firms. Overall, our evidence suggests that the pool of qualified female directors was sufficiently deep to avoid significant shareholder-borne costs of the quota.

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

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Valuation effects of Norway's board gender-quota law revisited*

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August 19, 2020

Abstract

We critically revisit and expand previous studies on the likely valuation effect of Norway's pioneering board gender-quota law. Most important, Ahern and Dittmar (2012) report a significantly negative average abnormal stock return, which they conclude is evidence of a large shareholder-borne cost of the quota constraint. We first show that they allocate their negative market reaction to the wrong event, which reverses their inference. We then show that their abnormal return estimate becomes statistically insignificant once we make the necessary adjustment for contemporaneous cross-correlation of stock returns. Furthermore, we provide new evidence on long-run abnormal stock returns, changes in Tobin's Q and operating profitability, and legal conversions, all of which corroborate that Norway's quota law most likely caused a statistically insignificant valuation effect on regulated firms. Overall, our evidence suggests that the pool of qualified female directors was sufficiently deep to avoid significant shareholder-borne costs of the quota.

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

1 Introduction

In December of 2005, the Norwegian government required public limited companies (ASA) to have gender-balanced boards within two years—or face forced liquidation. The board quota was driven by gender politics unrelated to firm performance (Government white paper *Odelstingsproposisjon* 97, 2002–2003), hence it provides a particularly interesting quasi-experimental setting for investigating the causal link between board gender composition and firm value as measured by the market reaction to quota news events.[1] Using this unique research setting, we critically revisit extant valuation evidence on Norway's quota law and show that the forced board gender-balancing did *not* impose significant costs on shareholders. Our analysis and new evidence are of general interest since several other countries are following Norway's lead and are adopting their own mandatory board gender quotas.[2]

There are two commonly discussed economic hypotheses in this context. The first is that restricting shareholders' free choice of directors results in lower board effectiveness. Firms may be forced to appoint female directors with less chief executive officer (CEO) experience, which may lower firm value (Fahlenbrach, Low, and Stulz, 2010; Ahern and Dittmar, 2012).[3] The second hypothesis holds that the quota increases the efficiency of board elections by reducing the influence of the male director "old boys" network (Agarwal, Qian, Reeb, and Sing, 2016) and increasing director independence and monitoring (Adams and Ferreira, 2009; Duchin, Matsusaka, and Ozbas, 2010; Masulis and Mobbs, 2011). Also, shareholders may benefit from the addition of a broader skill set when adding female directors (Kim and Starks, 2016; Bernile, Bhagwat, and Yonker, 2018; Adams, Akyol, and Verwijmeren, 2018).

Norway's pioneering quota law has attracted substantial empirical investigation of these two hypotheses. Bertrand, Black, Jensen, and Lleras-Muney (2019) document that the new female directors ushered

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

[2]Belgium, France, Germany, Iceland, Italy, the Netherlands, Portugal, and Spain, and the state of California have all recently passed gender balancing requirements for private firms. These countries typically impose substantially lower penalties for non-compliance than does Norway. See Kuzmina and Melentyeva (2020) and Hwang, Shivdasani, and Simintzi (2018) for studies of EU members and the state of California, respectively.

[3]Under this view, the cost of expanding board size—to make room for female directors while retaining male directors—places an upper bound on the expected shareholder-borne cost of the quota.

in by the quota were observably more qualified than their female predecessors in terms of education and professional experience, and have higher income. This suggests that the pool of qualified female directors may have been sufficiently deep to avoid significant shareholder-borne costs of the quota. Nevertheless, Ahern and Dittmar (2012) (henceforth AD) report a significantly negative market reaction to the initial quota announcement (in February of 2002), as well as a long-term decline in Tobin's Q of all-male boards through year-end 2009.[4]

The objective of this paper is to resolve the tension between the evidence of increased female director qualifications and AD's estimates of a large, negative stock market reaction to quota news for firms with all-male boards in 2001. We begin by revisiting AD's event-study. First, while they attribute the negative market reaction to a news event that *increased* the likelihood of the quota law, we show, using their own methodology, that the market reaction was statistically insignificant to this news event but significantly negative to a second event inside their five-day event window that *lowered* the quota likelihood. This evidence reverses AD's main conclusion. However, after correcting AD's standard errors for contemporaneous cross-correlation of stock returns, which they assume is zero, we can no longer reject the null hypothesis of zero abnormal stock returns to any of the two events. We follow up with a systematic examination of all the relevant quota news events (the relevant legislative period runs from 2002 through 2005), including firm-level cross-sectional regressions of abnormal stock returns. Overall, we cannot reject the hypothesis of a value-neutral market reaction to the quota news.

Following the analysis of short-term market reaction to quota news, we turn to long-run stock price performance over the period from February 2002 through April 2008 (when full compliance by all regulated firms had been achieved). We show that the portfolio of firms with all-male boards in 2001 and the portfolio of firms with at least one female director in 2001 both exhibit zero long-run stock return performance against standard risk-factor benchmarks. A long-short investment strategy in these two portfolios also exhibit a statistically insignificant long-run stock return performance. This exercise, which is similar in spirit to studies examining the performance of firms with 'high' and 'low' governance scores more generally (Gompers, Ishii, and Metrick, 2003), corroborates our short-term event-study evidence.

Next, we revisit AD's instrumental variable (IV) analysis, which they argue provide evidence of a large

[4] AD's conclusion received substantial attention in the financial press. For example, *The Financial Times* wrote "[Norway's quota caused] a large decline in Tobin's Q...over the following years." (August 20, 2011), while *The Economist* printed that "[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). Also, *The Wall Street Journal* wrote that "[the quota law] damaged shareholder value in the companies affected." (June 11, 2012).

(20%) quota-induced reduction in Tobin's Q of firms with all-male boards. AD instrument the annual board composition with the fraction of female directors in year 2002 interacted with year dummies until 2009. We argue that this instrument is endogenous and fails the exclusion restriction for an unbiased IV test. As AD shows, the fraction of female directors in 2002 is highly correlated with firm size, which itself directly affects Tobin's Q and hence arguably violates the exclusion restriction. Moreover, we provide new evidence in support of this critique.

We start by showing that AD's instrumentation has no significant effect on Q until the financial crisis—when it decreases Q. We then show that replacing AD's instrumentation with 2002 firm size (interacted with year dummies) generates similar IV regression results. This is consistent with a fundamental firm-size effect driving AD's results rather than the 2002 board gender composition. In another experiment, we exclude the five large government-controlled firms in AD's sample, which due to their superior access to capital performed relatively well during the financial crisis for reasons unrelated to their 2002 board composition (Beuselinck, Cao, Deloof, and Xia, 2017). This five-firm exclusion is sufficient to eliminate the statistical significance of the second-stage effect on Q. In sum, we conclude that AD's significant IV coefficient estimate has little to do with the quota constraint *per se*.

Finally, we address two pieces of evidence often used to support AD's finding of a negative valuation effect of the quota. First, while Matsa and Miller (2013) find that the ROA of ASA drops significantly relative to AS, we show that this ROA decline is most likely unrelated to the quota itself. Second, AD and Bøhren and Staubo (2014) claim that the quota caused a substantial number of ASA-to-AS conversions to avoid the quota constraint. However, after manually searching news and press releases, we show that no listed ASA delisted for reasons other than merger and acquisition (M&A) or bankruptcy—complex transactions that are almost certainly not driven by the quota restriction. For unlisted ASA, there are conversions throughout the entire sample period. However, we show that these unlisted ASA conversions are uncorrelated with the quota constraint after controlling for year fixed effects, which Bøhren and Staubo (2014) do not include. Year fixed effects are required to control for the strong quota-induced time-trend in the fraction of female directors.

The rest of the paper is organized as follows. Section 2 details the quota restriction, the legislative time line, and provides a sample description. In Section 3, we replicate AD's own abnormal return estimates as closely as possible and show how those estimates lack robustness. This is followed in Section 4 with our own multi-event abnormal return estimation, which fully incorporates cross-sectional dependence of stock returns, and cross-sectional regressions of firm-level abnormal returns. Section 5 examines long-term effects using portfolio stock returns and Tobin's Q. In Section 6, we present new and corroborating evidence on changes in operating profitability and legal conversions. Section 7 concludes the paper.

2 Quota restriction, time line, and sample characteristics

In this section, we first detail the quota restriction, which is necessary to define our central variable: the fraction of additional female directors required to comply with the quota ($Shortfall$). We then provide an overview of all of the main quota events that took place, which are new to the literature and subsequently used in our event-study (Section 4). Finally, we show characteristic of our sample of ASA firms.

2.1 The quota restriction

The quota applies to boards of ASA firms, but not the smaller and much bigger population of AS firms (ASA compares to the UK PLC, and AS compares to the UK Ltd). Under Norway's codetermination law, the shareholders elect one set of directors and the employees another set of directors (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 appointed for a term of two years. Overall, director elections are substantially influenced by shareholders, reflecting the fact that shareholder concentration is generally high (shown in Table 3 below).

As shown in columns (1) and (2) of Table 2, the quota mandates that, in a board with three directors, at least one must be female and at least one a male (the gender-balancing requirement). Moreover, there must be at least two women for boards with four to five members (the average board size among ASA), three women for six to eight-member boards, and four women for a nine-member board. Finally, for a board with ten or more members, the fraction of female (and male) directors must be at least 40%.

These restrictions imply that the fraction of female directors required to comply with the quota varies substantially with board size. To reflect this variation, columns (3)–(5) of Table 2 define $Shortfall_{it}$ as the difference in year t between firm i's fraction of female directors required by the quota and that of the current board. Throughout the paper, we use $Shortfall_{it}$ to measure the quota constraint facing an individual ASA in a given year t. Note that, since the required fraction of female directors varies with board size, the firm can affect $Shortfall_{it}$ either by replacing male directors or by changing board size.

2.2 Legislative time line and news events

After two government green papers (in 1999 and 2001) that considered various ways of increasing the number of women on corporate boards, a quota received unexpected government support in 2002. We focus our event study on the five news event dates listed in Table 1, 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, Ansgar Gabrielsen, 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—the Conservative Party ("Høyre")—which was the largest party in the coalition government. Likely as a result, Gabrielsen publicly retracted his support the very next day (not previously recognized by the literature). Over the following week, the parliamentary members of the Conservative party publicly reiterated the party's rejection of any quota imposed on private company boards. However, the smaller political party in the coalition government, the Christian Democratic Party ("Kristelig Folkeparti"), was open to the idea of a quota.

Given the clear stance of the Conservative Party against any quota, it came as a further surprise when, on International Women's day (March 8, 2002), the coalition government proposed a gender-quota law (event 2). The newspaper headline in *Dagens Næringsliv* that day reads (translated from Norwegian): "The Conservative Party outmanoeuvred." Taking a lead, the Cabinet promised compliance by government-owned firms within one year. The law proposal was submitted to Parliament on June 16, 2003 (event 3). It contained a provision stating that the quota law would be canceled if firms complied voluntarily by 2005. The 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 after reviewing voluntary compliance through 2005.

Although many firms immediately began to increase female director representation, the degree of voluntary compliance was ultimately deemed insufficient by the newly elected coalition government led by the Labour party ("Arbeiderpartiet"). (The Conservative party continued to fight against mandating the quota and, if the law were to be mandated, wanted to impose a soft sanction for non-compliance.) On December 1, 2005, the Prime Minister announced that the quota would be mandated and most likely include sanctions for non-compliance in the form of fines (event 4). However, when the Cabinet mandated the law on December 9 (event 5), the sanction effectively became forced liquidation—the ultimate penalty for violation of Norwegian corporate law. Firms were given two years to comply, and by April 2008 all firms subjected to the quota had complied.

2.3 Sample characteristics and board changes

We sample ASA from the *Brønnøysund Register Centre* database, 1998-2003, obtained through the database constructed by Berner, Mjøs, and Olving (2013). For a firm-year observation to be included, we require total assets>0, revenue>0, long-term assets≥ 0, current assets≥ 0, long-term debt≥ 0, short-term debt≥ 0, total assets>cash balances, and total assets>(current assets–current debt). Subsidiaries not reporting consolidated accounts are excluded. The total sample consists of 409 unique listed ASA and 888 unique unlisted ASA, of which 147 firms are listed in some years and unlisted in other years. Column (1) of Table 3 shows that, in a typical year, there are 174 listed ASA and 255 unlisted ASA. In terms of asset size (Column 3), listed ASA average more than three times that of unlisted ASA. Adding data from Norwegian tax authorities (2004-2013), Column (4) shows that stock ownership is on average highly concentrated and largely time-invariant: the largest shareholder has an average stake of 35% in listed ASA and 56% in unlisted ASA.[5] Column (5) of Table 3 and Figure 1 show that the fraction of female directors in all ASA rose from about 5% in 2001 to roughly 40% by early 2008. Moreover, Column (6) shows that the fraction of female chairs increased substantially from basically zero before 2001 to around 10% towards the end of the sample period.

As directors' 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), Column (8) describes board CEO experience. Notice first that, as shown in Column (7) and in Bertrand, Black, Jensen, and Lleras-Muney (2019), the fraction of female CEOs is generally low: in our sample it increases from 2% in 1998 to 5% for listed ASA and 10% for unlisted ASA in 2013. Column (8) shows the fraction of a board's directors with CEO experience.[6] For the average listed and unlisted ASA board, we see that board's CEO experience is stable over the sample period, at around 17%.

[5]Listed ASA are predominantly in Offshore/Shipping, Telecom/Technology, and Manufacturing, while unlisted ASA are in Financial Services and Telecom/Technology.

[6]We include CEO experience over the past three years in large firms: any ASA or a large AS, where the latter is among the 1% largest AS by revenue. We concentrate on CEO experience in the largest firms because the annual population of about 100,000 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 the 1% largest AS.

Turning to board size, Figure 1 shows that the average size of ASA boards remained at five shareholder-elected directors throughout the sample period. For a five-member board in 2003, quota compliance implied replacing 1.5 male directors with females, bringing female directors to two. Panel A of Figure 2 plots the ASA board-size frequency distribution in 2001 and 2008. It shows a narrowing of the distribution (significant at the 5% level with a non-parametric Kolmogorov–Smirnov test), reflecting a shift from four to five board members and from six to five. That is, while the quota did not cause a change in average board size, there is less variation around the average in 2008. This lower variation possibly reflects a desire to minimize quota-induced costs at the margin.[7]

Finally, Panel B of Figure 2 plots the frequency distribution of the number of board seats in ASA and the 1% largest AS held by male and female ASA directors in 2001 and 2008. As shown, ASA directorships are highly dispersed in both years: more than two-thirds of individual directors hold only one board seat. Moreover, the distribution is largely similar for male and female directors. Consistent with Bertrand, Black, Jensen, and Lleras-Muney (2019), Panel B suggests that the pool of qualified female directors was sufficiently large to prevent a disproportionally high seat concentration among females directors.

3 Revisiting AD's event-study

In this section, we critically revisit AD's main conclusion that the first event listed in Table 1 (Friday, February 22, 2002) caused a significantly negative valuation effect for OSE-listed companies with all-male boards. We first replicate as closely as possible AD's abnormal return estimates using their sample of firms, return data, and econometric methodology. We then bring two important missing components into AD's analysis, which uses the five-day window (-2,2) centered on Friday, February 22. The first is a second major news announcement on Saturday, February 23, 2002, which is ignored by AD and which *reduced* the likelihood of a quota. The second missing component in AD's analysis is the contemporaneous cross-correlation of stock returns, which exists when the event affects all firms simultaneously in calendar times.

[7] As shown in Table 2, increasing the board from four to five members allows a firm to appoint two females while retaining three (rather than two) males, and reducing board size from six to five directors allows the firm to appoint two (rather than three) females while retaining three males.

3.1 A close replication of AD's event study

Panel A of Table 5 copies AD's average five-day abnormal stock returns directly from their Table III. For each of their total sample of 94 OSE-listed ASA i, the abnormal return is computed as follows:

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

where r_i is firm i's return 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. Day 0 is Friday, February 22, 2002, when the Minister of Industry and Trade Gabrielsen surprisingly announced his support of a board gender quota (the first event in Table 1 above). AD's stock return data are from Compustat Global for the listed ASA and from the Center for Research in Securities Prices (CRSP) for the industry-matched US firms. AD classifies the sample firms based on their board composition in year 2001—prior to the February 2002 news event—obtained from firms' annual reports.

In Panel A of Table 5, the average $CAR_i(-2,2)$ is -2.57%, which has a p-value of 0.001 and thus is significantly negative at the 1% level. The p-value is computed based on the cross-sectional standard deviation of $CAR_i(-2,2)$. Furthermore, columns (2)–(4) show that the negative market reaction is concentrated among firms with zero female directors in 2001 ($Zero_{2001}$). In Column (4), the difference in average abnormal return between the firms with all-male boards and firms with at least one female director (Pos_{2001}) is -3.52%, which is highly significant. Based on this evidence, AD conclude that the average market reaction to the February 22, 2002, quota event, which increased the likelihood of a future quota law, was significantly negative.

We attempt to replicate Panel A using information available in AD. This includes Eq. (1), the names of their 94 sample firms and, as closely as possible, their return data from Compustat Global and CRSP.[8] As AD do not provide complete information on their data selection criteria, we restrict the matching firms to NYSE/Amex/Nasdaq common stocks (CRSP share codes 10 and 11) of firms incorporated in the US. Moreover, in the presence of multiple share classes, we select the share class with the highest trading volume over the estimation period. Most important, as AD provide no information on how they treat non-trading days inside the five-day event window, we adopt two alternative return assumptions in

[8] We thank Kenneth Ahern for providing the names of the 94 firms. While AD's board data are from annual reports, ours come from *Brønnøysund Register Centre*, where 69 firms (68 in AD) have zero female directors in 2001. This one-firm difference does not affect the statistical inferences below.

panels B and C.

First, in Panel B (which contains two sub-panels B1 and B2) of Table 5, we require at least one trade within the five-day event window. As it turns out, it is necessary to include multi-day returns—some of which extend prior to day -2—to generate a $CAR(-2, 2)$ estimate for each of AD's 94 sample firms. As shown in Panel B1, the average $CAR(-2, 2)$ of -2.83% is statistically significant at the 1% level when we, like AD, assume cross-sectionally independent returns. For the 69 firms with all-male boards ($Zero_{2001}$) in Column (2), the average $CAR(-2, 2)$ is -3.89% and, as in Panel A, significantly different from that of firms with at least one female director (Pos_{2001}) in Column (3). Thus, when requiring at least one trade within the five-day event window, we are able to include all 94 firms in AD and reach the same statistical inference as AD. In Panel C, we modify the return requirement in Panel B by requiring trades on two consecutive days between day -3 and +2. This restricts the sample available for estimating AD's $CAR(-2, 2)$ from 94 to 79 firms (and to 65 firms if we instead require an exact five-day return). For this subsample, average $CAR(-2, 2)$ is -3.26% for $Zero_{2001}$, and 0.11% for Pos_{2001}. Moreover, as shown in Column (4), the difference in the mean is significant at the 10% level only.

As pointed out by AD as well, the percent of female directors in 2001 is highly correlated with firm size. Note also that AD's 94 OSE-listed sample firms are a magnitude smaller than the US matching firms in Eq. (1)—they are not matched on firm size, whether belonging to $Zero_{2001}$ or Pos_{2001}. It is therefore possible that AD's difference in average CAR across $Zero_{2001}$ or Pos_{2001} is driven by differences in OSE-listed firm size rather than the quota constraint. We examine this possibility in Panel B2 of Table 5. While maintaining the sample sizes in B1, B2 redefines (in this panel only) $Zero_{2001}$ (Pos_{2001}) to be the 69 smallest (25 largest) sample firms, measured by total revenue in 2001. The average $CAR_i(-2, 2)$ of the redefined $Zero_{2001}$ and Pos_{2001} are now -3.66% and -0.53%, respectively, which are significantly different at the 10% level in Column (4). The similarity of the abnormal returns in panels B1 and B2 suggests that AD's abnormal return difference in Column (4) may reflect differences in firms size. We return to the potential role of firm size when discussing AD's analysis of Tobin's Q in Section 5.2.

3.2 Adjusting for the second event inside AD's event window

In Panel D of Table 5, we introduce a day-by-day analysis of AD's five-day event window. Although AD did not report day-by-day results, this allows us to examine whether AD's large negative average $CAR(-2, 2)$ is actually driven by the February 22 news event, as they assume, or the next day's *probability-decreasing* event. To be included, a firm must have a one-day return on the event day in question, which produces 64–67 abnormal return (AR) observations per day. We continue to use AD's method for computing p-values (i.e., assuming cross-sectional independence of the individual AR_i). These assumptions ensure that AD could have performed the same day-by-day analysis had they so chosen.

Recall from Table 1 that the Minister of Industry and Trade on Saturday, February 23, 2002, *publicly retracted* his support for the quota expressed one day earlier. His retraction, which appeared in Norway's major national business daily *Dagens Næringsliv*, was consistent with his own ruling-party's negative view on a board gender-quota law (the newspaper headline reads (translated from Norwegian) "Gabrielsen no longer supports a quota"). Thus, while his Friday support statement likely increased the market's assessment of the probability of a quota law, his next-day (Saturday) retraction of the support must have reversed this probability assessment. The market reaction to the Saturday news event materialized on the next trading day, Monday, February 25.[9]

Since the Monday is day +1 in AD's five-day event window, one cannot infer from AD's results whether the negative average $CAR(-2, 2)$ represents the market reaction to the first or the second event. If the negative reaction occurs on Friday and not on Monday, the conclusion is that the market was negative to the prospect of a new quota law. However, if the negative market reaction occurs on Monday and not on Friday, the opposite conclusion holds—that the market was positive to the prospect of a quota law. We separate each day in AD's five-day event window in order to find the answer to this crucial question.

In panel D of Table 5, we denote the one-day average abnormal return estimate as $AR(\tau)$, where $\tau = 1, ..., 5$ is the day inside the five-day event window. Note first the statistically insignificant AR on Friday, February 22. That is, contrary to AD's inference, the average market reaction to the news that increased the likelihood of a future quota law was *not* significantly different from zero even under AD's assumption of zero cross-correlation of returns.[10] However, the average market reaction on Monday, following Saturday's probability-reducing quota news, was *negative* and statistically significant, with $AR = -1.99\%$. Moreover, as shown in columns (2)–(4), this negative market reaction is concentrated among firms with all-male boards. In sum, AD assign their own evidence of significantly negative abnormal returns to the wrong

[9]On p.155, AD states: "The public announcement [on February 22, 2002] was the top story in Norway's largest newspaper, Verdens Gang (VG) with the headline (translated from Norwegian) 'Sick and Tired of the Old Men's Club!"'. It is worth pointing out, however, that while VG is a leading national 'gossip' newspaper, *Dagens Næringsliv*, in which the Minister of Industry and Trade found it necessary to retract his Friday message the very next day, is Norway's leading national business newspaper.

[10]Neither a t-test for difference in mean nor a Wilcoxon rank-sum test for difference in median indicate that the quota-announcement on Friday, February 22, 2002, affected firms labelled $Zero_{2001}$ or Pos_{2001} differently.

event. Using their own data and methodology, the correct conclusion is that the market reacted negatively to news that a quota law would be *less likely* than previously anticipated. This reversal fails to support the hypothesis that the market considered the quota law to be costly for shareholders, and even suggests the opposite market reaction.

We next turn to our second critical assessment of AD's event-study methodology: p-values that are biased downward due to their assumption of cross-sectionally independent abnormal stock returns.

3.3 Adjusting AD's p-values for return cross-correlation

Stock returns tend to move in the same direction in calendar time regardless of any heterogeneous firm-level impact of a news announcement. Therefore, AD's assumption that the individual returns, $CAR_i(-2,2),\ i = 1, ..., 94$, are cross-sectionally independent over the five-day calendar-time period biases downward their p-value of the average $CAR_i(-2,2)$. To illustrate the magnitude of this bias, let all firms have identical individual standard deviation of daily returns (σ) and pairwise daily return correlation (ρ). The standard deviation of the average return across N firms is (Kothari and Warner, 2007):

$$\sigma_N(\rho) = \sqrt{\frac{1}{N}\sigma^2 + \frac{N-1}{N}\sigma^2\rho}, \tag{2}$$

and the bias from assuming $\rho = 0$ is therefore

$$\frac{\sigma_N(\rho)}{\sigma_N(\rho = 0)} = \frac{\sigma_N(\rho)}{\sqrt{\sigma^2/N}} = \sqrt{1 + (N-1)\rho}. \tag{3}$$

For example, with $\rho = 0.10$, which is a typical average pairwise return correlation on the OSE (Næs, Skjeltorp, and Ødegaard, 2008) and in US stock markets (de Bodt, Eckbo, and Roll, 2020), AD's sample size of $N = 94$ means that $\sigma_N(\rho) = 3.2\sigma_N(\rho = 0)$. In other words, the true standard deviation is 3.2 times greater than a standard deviation assuming zero contemporaneous cross-correlation of returns. Interestingly, this simple illustration is close to the empirical bias in AD's p-values identified below.

To properly adjust AD's significance levels for $\sigma_N(\rho)$, it is necessary to introduce a time-series analysis into the event study. We therefore run the following portfolio time-series regression:

$$r_{pt}^{-I} = \alpha_p + AR_p d_t + \epsilon_{pt}, \tag{4}$$

where $r_{pt}^{-I} \equiv \frac{1}{N}\sum_{i=1}^{N}(r_i - r_{i,match})_t$ is the equal-weighted average daily abnormal return using AD's definition in Eq. (1). The estimation period is day -252 through day 2, where day 0 is February 22, 2002, and it uses daily stock returns from Compustat Global and CRSP. The dummy variable d_t takes a value of one during AD's event window (-2,2) and zero otherwise. Thus, AR_p is simply the daily average abnormal stock return over the five-day event window and $5AR_p$ is the average $CAR_i(-2,2)$. Moreover, as the portfolio return r_{pt}^{-I} fully absorbs whatever contemporaneous cross-correlation exists at time t, the standard error of AR_p is $\sigma_{AR}(\rho)$, which corresponds to $\sigma_N(\rho)$ in the illustration above. We use the correct t-statistic $5AR_p/\sigma_{5AR_p}(\rho) = AR_p/\sigma_{AR_p}(\rho)$ to test the null-hypothesis that AD's average abnormal return $CAR(-2,2)$ is equal to zero.[11]

Recall from Panel C of Table 5 that, when we restrict each firm to have at least a single one-day return in the event window (-2,2), AD's sample is reduced from 94 to 79. This is the base-sample for the results reported in Panel A of Table 6. We again start with AD's five-day window, but this time use the time-series portfolio regression in Eq. (4) to incorporate the cross-sectional correlation. The average $CAR(-2,2) = 5AR_p$ is -2.20% and, as expected, almost identical to the average $CAR(-2,2)$ reported in Panel C of Table 5. However, the p-values are now much higher, to the point where one can no longer reject the hypothesis of $AR_p = 0$ even for the subsample of all-male boards.

Panel B reports the day-by-day portfolio estimates of abnormal returns, again requiring a one-day return in the event window. Importantly, the abnormal return on Monday, February 25, 2002, is now also insignificantly different from zero, as are all individual days in AD's five-day event window.[12] In sum, AD's own conclusion of a statistically significant negative market reaction to the quota announcement, as well as the significance of Monday, February 25, in our expanded day-by-day analysis, all require the untenable assumption of a zero contemporaneous cross-correlation of stock returns.

In sum, in this section we have shown that AD's negative abnormal return estimate for firms with all-male boards is driven by the second (Saturday's) probability-decreasing announcement and *not* by the Friday's probability-increasing announcement, which effectively reverses their main conclusion. We also

[11]While not tabulated, we have verified that our results are robust to estimating the abnormal returns using a system of seemingly unrelated regressions (SUR), with a single OLS regression for each sample firm, and computing standard errors accounting for the residual cross-correlation (Kolari and Pynnönnen, 2010). 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).

[12]While not tabulated, this conclusion holds also if we include a dummy variable in Eq. (4) for day t being a Monday in the estimation period. This confirms that there is no so-called weekend effect (French, 1980) driving our abnormal return estimates.

show that adjusting for the cross-correlation of stock returns created by an event that affects all firms simultaneously in calendar time results in p-values that are too large to reject the null hypothesis of zero abnormal returns for *any* of the days inside AD's five-day event window—otherwise estimated using AD's abnormal return definition.

4 A comprehensive multi-event analysis

In this section, we perform a comprehensive analysis of the market reaction to all five quota events listed in Table 1. We first introduce the portfolio-based approach, which is required to control for any contemporaneous cross-correlation of stock returns, and then provide cross-sectional estimates of the determinants of firm-level abnormal returns.

4.1 The portfolio-based approach

To maximize test power (Brown and Warner, 1980, 1985), we use the standard two-day event window (-1,0), which ends with the public announcement date (day 0). Note that this two-day window allows us to estimate the market reaction to February 22, 2002, without it being confounded by the second announcement on event day +1 (February 25). For the same reason, we do not include February 25 here as its day -1 would be February 22 (the market reaction on February 25 is already properly estimated and reported in Panel B of Table 6).

For each of the five events, we estimate $CAR(-1, 0) = 2AR_p$ for a portfolio of OSE-listed firms using the return-generating process:

$$r^e_{pt} = \alpha_p + \beta_p r^e_{wt} + AR_p d_t + \varepsilon_{pt}, \tag{5}$$

where r^e_{pt} 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_t is a dummy for the event window (-1,0), and r_{wt} is the daily excess return on the MSCI stock market world index. The regression starts 252 trading days prior to and ends on event day 0, excluding days of prior events, if any. To be included in portfolio p, a firm must have a minimum of 100 one-day return observations from *Oslo Børsinformasjon* and a one-day return observation on each day in the two-day event window. This leads to portfolios with the number of listed firms ranging from 132 to 146 across the five events.[13]

[13]Daily stock returns from *Oslo Børsinformasjon* are computed using differences in (log of) daily closing prices, adjusted

In Table 7, *High shortfall* firms have a female director $Shortfall$ at or above the median in the preceding year-end. Columns (1)–(4) use samples of OSE-listed Norwegian ASA, while columns (5)–(7) use samples of OSE-listed Norwegian ASA (treated) and OSE-listed foreign-domiciled (control) firms in the oil/offshore sector. At the bottom of the table, we re-estimate the model across all five events simultaneously, with the dummy variable d taking a value of one in the event window (-1,0) for all five events, which produces $CAR_{1-5}(-1,0) = 10AR_p$. This last estimation begins 252 days prior to the first event (February 22, 2002) and ends on the day of the fifth event (December 9, 2005). This particular estimation is added as an overall check on the sum of the market reactions to the five events.

With the exception of event 2 (March 8, 2002), none of the five news events in Table 7 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 events* using $CAR_{1-5}(-1,0)$.[14] As to the positive $CAR(-1,0)$ on March 8, 2002, 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.[15] Note that Column (5) reports a weakly positive $CAR(-1,0)$ the 31 OSE-listed ASA operating in the oil/offshore sector, while 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 on March 8, 2002, appear to be industry-driven, rather than a phenomenon tied to Norwegian incorporation and hence the quota law.

4.2 Cross-sectional determinants of abnormal returns

In this section, 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 as well as other

for splits and dividends. In the estimation period, if a closing price is missing, it is replaced by the bid-ask midpoint if available. Twenty percent of the returns in the estimation period are generated from bid-ask prices.

[14]While not tabulated, this conclusion holds with an alternative 3-day event window (-1,1), irrespective of the risk adjustment. Moreover, our inferences are unaffected if we instead 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).

[15]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.

firm-specific characteristics. 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:

$$CAR_{ik}(-1,0) = \alpha_k + \gamma_{1k} Shortfall_{ik} + \gamma_{2k}\mathbf{X_{ik}} + \kappa_i + u_{ik}. \quad k = 1, \ldots, 5. \tag{6}$$

In addition to $Shortfall$, the vector of controls $\mathbf{X}$ includes several firm characteristics. *Largest owner* is the percent ownership of the largest shareholder. *Government control* is a dummy variable indicating a government ownership of at least 30% of the outstanding shares. *Codetermination* is a dummy indicating that quota-induced females and employee directors together have a majority of the board seats. *Risk* is the firm's daily stock return volatility in the year prior to the event and *Total assets* is log of book value of total assets. All variables are defined in Table 4.

As is commonly known, board characteristics are endogenous and correlated with firm characteristics. Hence, since $Shortfall$ is correlated with firm size, we include *Total assets* to help control for size effects on board composition that is unrelated to the quota itself. *Largest owner* and *Government control* capture large owners that have considerable influence over director appointments. *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 results are shown in Table 8. Importantly, the regressions fail to identify significant effects of $Shortfall$ on the event returns for all five events.

In sum, the average abnormal return estimates in Table 7, which account for the contemporaneous cross-correlation of stock returns, fail to reject the null hypothesis of a zero two-day market reaction to the gender-quota news announcement in Table 1, whether individually or jointly. Moreover, the cross-sectional regression coefficient estimates in Table 8 fail to reject that the individual firm-level abnormal stock returns to any of the announcements are correlated with the fraction of female directors.[16]

[16] This conclusion is robust also to using $Shortfall_{number}$—the number instead of the fraction of female directors—in the empirical analysis.

5 Long-run performance estimates

Under market efficiency, the average market reaction to the quota news announcement—estimated in Section 3 above—represents an unbiased estimate of the true valuation effect of the quota constraint. Therefore, if the valuation effect is truly close to zero—as suggested by our evidence of statistically insignificant market reaction—there should be no subsequent (long-run) quota-induced abnormal performance of the OSE-listed firms. In this section, we first test this proposition conditional on a model generating expected returns. This analysis, which is similar in spirit to extant studies of the long-run performance of stock portfolios of 'high' versus 'low' governance quality (Gompers, Ishii, and Metrick, 2003), therefore serves as a useful check of the main statistical inference in Section 4 above. We then turn to AD's time-series analysis of Tobin's Q and critically examine their inference, which is at odds with our evidence of statistically insignificant event-induced abnormal stock returns.

5.1 Long-run portfolio performance

We measure long-run abnormal portfolio performance using the parameter α_p in the following three-factor model:

$$r^e_{pt} = \alpha_p + \beta_{p1} r^e_{wt} + \beta_{p2} HML_t + \beta_{p3} SMB_t + \varepsilon_{pt}, \quad t = 2/2002, ..., 4/2008. \tag{7}$$

where r^e_{pt} now is the monthly USD-denominated stock return to portfolio p of domestic OSE-listed ASA, which is converted to USD using the monthly exchange rate and in excess of the current month's 3-month US Treasury bill. r^e_{Wt} is the monthly return on MSCI world stock market index in excess of the current month's 3-month US Treasury bill. HML_t and SMB_t are monthly returns to the global value and size factors (Fama and French, 1993) found on Kenneth French's web site. This portfolio estimation uses monthly portfolio returns beginning in February 2002—the first major quota event—and ending in April 2008 when all ASA were in compliance.

Table 9 shows α_p estimates for three alternative equal-weighted portfolios. In columns (1)-(3), the return generating process is Eq. (7) above, while columns (4)-(6) add a global momentum risk factor, MOM (Carhart, 1997), also found on Kenneth French's web site. The first portfolio, $Zero_{2001}$, contains an average of 98 OSE-listed ASA with 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_p is insignificantly different from zero for all three portfolios. That is, even for a portfolio that goes long in firms that are the most affected by the quota ($Zero_{2001}$) and short in the least affected firms (Pos_{2001}), there is no long-run abnormal stock performance. This evidence supports our conclusion from the short-run event study of a value-neutral market reaction to forced board gender-balancing.

5.2 A critique of AD's IV test for effects on Tobin's Q

AD examine long-term effects of the quota on Tobin's Q using a two-stage IV regression, with reference to Stevenson (2010)'s study of the effect of female athletic participation on their eduction and professional outcomes. AD's IV analysis is designed to account for the effect of firms' endogenous quota compliance timing over the six-year period, 2003–2009. To achieve this, the base for their instrumentation in the first step is the 2002 fraction of female directors, which they interact with year dummies, 2004–2009. However, as discussed below, this instrumentation most likely fails the exclusion restriction and hence is invalid in AD's context.

To illustrate AD's IV procedure, we employ our variable *Shortfall*, which corresponds to AD's fraction of female directors but will produce an opposite coefficient sign in a regression on Tobin's Q. AD's first step regression can then be written as follows:

$$Shortfall_{it} = \alpha + Shortfall_{i,2002} \sum_{t=2004}^{2009} \beta_t \tau_t + \theta_i + \tau_t + u_{it}, \quad t = 2003, ..., 2009, \tag{8}$$

where θ_i and τ_t are industry and year fixed effects, respectively, and the summation is zero in year 2003. The second-step regression is

$$Q_{it} = \alpha + \beta \widehat{Shortfall}_{it} + \theta_i + \tau_t + \epsilon_{it}, \quad t = 2003, ..., 2009, \tag{9}$$

where $\widehat{Shortfall}_{it}$ is the predicted value from the first step. We use *Oslo Børsinformasjon* to measure Q (see Table 4) for a definition.[17] As shown in Column (1) and Panel A of Table 10, our estimation of Eq (9) results in a coefficient estimate of 1.91 on $\widehat{Shortfall}_{it}$, which is statistically significant at the 5% level and almost identical to AD's estimate of -1.94. The change in sign is, of course, due to the fact

[17] We eliminate firm-years with Q$\leq$ 0, and winsorize the remaining observations at 1% and 99% each year. While AD employ industry-adjusted Q, we follow the recommendation of Gormley and Matsa (2014) and exclude the industry adjustment. Our total number of firm-years is 820 (compared to 630 in AD), reflecting our access to board composition data in *Brønnøysund Register Centre* and stock-return data from *Oslo Børsinformasjon*.

that AD use the fraction of female director, whereas we use the inverse shortfall of female directors. This coefficient estimate is what prompts AD to conclude that "[Norway's gender quota] led to value losses of upwards of 20% for the firms with [all-male boards]" (p. 168).[18]

Column (1) in Panel B of Table 10 shows that the instrument $Z_{it} \equiv Shortfall_{i,2002} * \tau_t$ ($t = 2004, ..., 2009$) is a good predictor of $Shortfall_{it}$. Formally, the hypothesis of joint instrument irrelevance across years, i.e, that $Cov(Z_{it}, Shortfall_{it}) = 0$ for all t=2004,...,2009, is strongly rejected by the F-value of 85.9. Therefore, the instrumentation meets the inclusion restriction (Angrist and Pischke, 2009). However, it is difficult to argue in economic terms that the instrument also meets the exclusion restriction. This restriction requires Z_{it} to impact Q_{it} only through $\widehat{Shortfall}_{it}$, so that $Cov(Z_{it}, \epsilon_{it}) = 0$. This condition is unlikely to be met since board characteristics are endogenous and correlated with firm characteristics, which in turn affect Q. Therefore, Z_{it} is most likely correlated with the error term ϵ_{it} in Eq. (9). Note also that, because the instrument Z_{it}, by construction, has a time-varying effect on Q, this correlation is not eliminated by AD's inclusion of year- and firm fixed effects in Eq. (9). Hence, we argue that the endogeneity bias caused by latent time-varying omitted variables is not eliminated and the exclusion restriction is violated.

Table 10 provides new evidence that support the above argument. Consider first the impact of $Shortfall_{2002} * \tau_t$ in the reduced-form regression estimation in Column (2) of Panel B. This reduced-form regression has identical regressors as Eq. (8) but uses Q_{it} as dependent variable. Note that $Shortfall_{i,2002} * \tau_{2008}$ is the only statistically significant coefficient estimate. That is, it appears that the presumed effect of the 2002 board composition on Q does not appear until 2008. A similar conclusion emerges from AD's reduced-form regression (Panel B of their Table IV). They explain this late impact of $Shortfall_{i,2002}$ on Q with firms' last-minute quota compliance (in 2007).

However, as discussed above, since the fraction of female directors in 2002 is endogenous and correlated with firm size, instrument exogeneity is likely violated ($Cov(Z_{it}, \epsilon_{it}) \neq 0$). To examine whether the IV result of AD is driven by firm size rather than board composition *per se*, Column (3) shows the result of replacing $Shortfall_{2002}$ with $log(Sales_{2002})$ in the instrumentation. Notice first that $log(Sales_{2002}) * \tau_t$ also satisfies the relevance condition (F-value of 44.1 in Panel B). More importantly, the second-stage coefficient estimate is highly significant (Panel A), which suggests that larger firms in 2002 performed

[18]Beginning in 2005, listed firms were required to report using the International Financial Reporting Standard (IFRS), switching from Norwegian General Accepted Accounting Principles (NGAP). While not tabulated, adding three firm characteristics from Table 4—*Total assets*, *Leverage*, and *Board size*—as regressors in the IV test does not change our inferences.

better than smaller firms, in particular during the period 2007–2009 (Column 4 of Panel B). This challenges AD's interpretation that the fraction of female directors in 2002 is the fundamental driver of Q. It is more likely that the fundamental driver is firm size, which correlates with the 2002 board gender composition.

Moreover, we suggest an alternative explanation for the late impact of $Shortfall_{i,2002}$ on Q that recognizes firms' differential responses to the financial crisis. To illustrate, note that four of the five government-controlled sample firms fully complied with the proposed quota already in 2002 (and the fifth in 2003). Moreover, as their government connection gave these firms superior access to capital, they performed relatively well during the financial crisis for reasons other than their 2002 board composition (Beuselinck, Cao, Deloof, and Xia, 2017). This suggests that a proper test of AD's proposition should exclude these five firms. As shown in columns (5) and (6), when we exclude the five government-controlled firms, the coefficient estimate for $\widehat{Shortfall}$ in the second IV stage becomes statistically insignificant, as does the coefficient estimate for $Shortfall_{2002} * \tau_{2008}$ in the reduced-form regression. In other words, the shortfall of female directors has no statistical impact on Q once the five government-controlled firms are excluded from AD's IV regression.

Finally, we highlight yet another weakness of AD's IV analysis. Recall from Table 5 that AD correctly classify their sample firms' board composition in year 2001 as exogenous to the market reaction to the February 22, 2002, quota-news event. In contrast, in their IV analysis, they instead use the board composition at year-end 2002 as providing an exogenous cross-sectional distribution of the fraction of female directors.[19] This is problematic because, following the March 8, 2002, announcement, as much as 29% of the OSE-listed firms reduced $Shortfall$ at the annual shareholder meetings in the spring of 2002.[20] As shown in Column (7) of Table 10, replacing $Shortfall_{2002}$ with $Shortfall_{2001}$ indeed results in a coefficient estimate on $\widehat{Shortfall}$ of 0.69, which is statistically insignificant.

Overall, the Q-analysis in Table 10 casts serious doubt on the validity of AD's inference from their IV test, and it corroborates our main conclusion of a largely value-neutral effect of the quota law.[21]

[19] "We use the 2002 annual reports to measure exogenous variation in the mandated board change, a full year before the quota was passed in December 2003" AD, p.161.

[20] Recognizing this endogeneity, AD's statement that "[t]o verify that the gender of the boards was not yet impacted in 2002, we compare the gender composition of the boards in 2001 to 2002 and find that the majority of the firms had the same gender composition in both years." (p.161) does little to resolve this obvious problem of endogeneity.

[21] We have also verified that replacing $Shortfall_{2002}$ with the number of missing female directors under the quota requirement fails to produce a statistically significant coefficient estimate in the second-stage IV regression. Also, while finding an IV instrument that works in AD's setting goes beyond the purpose of this paper, we experimented with the instrument used by Adams and Ferreira (2009) in a different context: the fraction of male directors on the board who sits on other boards

6 Additional corroborating evidence

The above evidence suggests that the Norwegian gender quota imposed few if any costs on shareholders of OSE-listed ASA. In this section, we broaden the debate by revisiting the conclusion of two additional studies which, like AD, argue that the quota imposed large costs on ASA. The first is Matsa and Miller (2013), who report estimates of a significant decline in the post-quota operating performance of ASA relative to AS. The second study is Bøhren and Staubo (2014), who conclude that many ASA switched legal form to AS in order to escape a costly quota constraint.

6.1 Changes in operating profitability

Using a diff-in-diff regression over the period 2003–2009, Matsa and Miller (2013) report a post-2006 decline in ROA (defined as EBIT/total assets, see Table 4) for listed ASA relative to a sample of matched AS. While they also examine an alternative control group of Nordic listed companies, we focus on their use of Norwegian AS. They estimate the quota-effect on ROA using the following type of OLS regression:

$$ROA_{it} = \gamma_0 + \gamma_1 ASA_i * Comply_t + \gamma_2 \mathbf{X_{it}} + \theta_i + \tau_t + \epsilon_{it}, \tag{10}$$

where $Comply_t$ takes a value of one for the period 2007–2009, and $\mathbf{X_{it}}$ is a vector of firm characteristics. Their coefficient estimate on $ASA * Comply$ is –0.027 (columns 1 and 2 of their Table 3). They conclude that this reduction in short-term profits in the post-quota period is because listed ASA "undertake fewer workforce reductions than comparison firms, increasing relative labor costs and employment levels" (abstract).

Matsa and Miller (2013)'s diff-in-diff analysis (Eq. 10) assumes that listed ASA and AS exhibit otherwise identical responses to aggregate shocks. This assumption is questionable since the choice of being a listed ASA or an AS is endogenous and may be correlated with latent time-variant factors not captured by the fixed effects. While we do not attempt to resolve this concern, we revisit Eq (10) after broadening the treatment group to include unlisted ASA. Also, we use a longer time period to address whether the negative ROA effect identified by Matsa and Miller (2013) lasts beyond the financial crisis period.

on which there are female directors. In our setting, this instrument fails the inclusion restriction (i.e., is uncorrelated with $Shortfall$) once we add year fixed effects to account for the time trend caused by quota compliance.

Our results of estimating Eq. (10) are shown in Table 11. The indicator variable $Comply_t$ is one for years $t \geq 2008$, as compliance was required by 12/2007, and zero otherwise. The vector $\mathbf{X}_{it}$ contains *Firm age*, *Size*, *Leverage*, *Largest owner*, *Board CEO experience*, *Board size*, and *Board busyness*, as defined in Table 4 above. In order to produce a control sample of AS of comparable firm size to ASA, we use the 1% largest AS by revenue.[22] We exclude financial firms and ASA that were registered as AS at any time during the sample period, and require firms to have at least two observations and no missing values for the control variables. The total sample consists of 436 ASA and 1,786 large AS.

In columns (1) and (2) of Table 11, we use the 2003–2009 sample period in Matsa and Miller (2013). Much as in their paper, this produces a coefficient estimate for $ASA * Comply$ of –0.024 and –0.022, both significant at the 5% level. However, when we extend the sample period through 2013, the interaction variable $ASA * Comply$ becomes statistically insignificant (columns 3 and 4). Moreover, when we decompose $Comply_{it}$ into year-by-year effects ($ASA_i * \tau_t$), columns (5) and (6) show that there is a negative effect on ROA in 2008 only and a significantly *positive* effect in 2013 after inclusion of firm characteristics.

In sum, Table 11 suggests that the negative treatment effect identified by Matsa and Miller (2013) is limited to year 2008. As such, it may well be the result of a heterogeneous impact of the financial crisis on treated and control firms as opposed to tracing back to the quota constraint.[23]

6.2 Legal conversions

Recall that the quota law applies to ASA only, and a firm must be ASA to be listed on the OSE. AD and Bøhren and Staubo (2014) conclude that OSE-listed ASA left the stock exchange to avoid the quota constraint. In contrast, we show below that *no listed* ASA converted to AS for reasons other than being the target in an M&A transaction or liquidation in bankruptcy. Moreover, we show that the propensity of unlisted ASA to convert to AS is uncorrelated with the female director shortfall in the cross-section.

Note first that firms never publicly explain a conversion in terms of the quota law, and so the researcher

[22] The large AS have an average asset size that is similar to that of unlisted ASA and a revenue size that lies between that of listed and unlisted ASA. In terms of board gender composition, the gender quota had a limited spill-over effect on the boards of large AS, where the fraction of female directors increased from 8% in 2001 to 13% in 2013. This modest increase reflects a general trend throughout western economies over the sample period. 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*).

[23] Our results are robust to using, as do Matsa and Miller (2013), a matched sample of AS as control group. 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."

must decide which observed conversions are reasonably driven by the quota constraint. AD choose to characterize a delisting following an M&A transaction as driven by quota avoidance. In contrast, we exclude M&As from the conversion analysis because these costly transactions are primarily driven by the synergy gains from merging the two firms (Betton, Eckbo, and Thorburn, 2008; Eckbo, 2014). Moreover, keep in mind that a delisting just to avoid the quota constraint is costly due to the loss of listing benefits, which include liquidity and access to public equity. For unlisted ASA, however, converting to AS is low-cost because an unlisted ASA face near-identical regulations as an AS in terms of accounting and governance standards (Bøhren and Staubo, 2014).[24]

Panel A of Figure 3 shows the annual number of non-financial listed ASA, delistings due to M&A and bankruptcy, and new lists, 2002–2009.[25] Financial firms are excluded because, until 2007, these firms were required by law to be ASA. Our conversion data from *Brønnøysund Register Centre* are complemented by a search of news and press releases for M&As involving a foreign acquirer. There are two main findings related to listed ASA. First, as expected and shown in Panel A, not a single listed ASA delisted for a reason other than an M&A transaction or bankruptcy. Second, the total number of listed firms increased steadily over the sample period.[26]

Turning to Panel B of Figure 3, there was a steady decline in the total number of unlisted ASA between 2001 and 2009. A majority (61%) of the firms leaving the unlisted-ASA category were acquired in an M&A transaction, liquidated in bankruptcy, or went public (became listed ASA). This leaves a residual group of 165 ASA-to-AS conversions for reasons that are unknown. These conversions are therefore candidates for empirically testing whether they were driven by the quota constraint *per se*. We use the following logit model to estimate the likelihood that these "unknown" conversions are correlated with the female director shortfall:

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

As before, κ_i and τ_t are, respectively, industry and year fixed effects, and the vector $\mathbf{X}_{it}$ of explanatory variables are as in Table 11. The dependent variable, $Convert_{it}$, equals one if the firm converts to AS

[24]For an unlisted ASA that abandons plans to list on the OSE, there are few benefits of remaining an ASA while there are some extra costs relative to an AS. These costs include having to register the shares with Norway's securities registry (VPS) and some additional corporate governance restrictions.

[25]In 16% of the acquisitions the firm continued as an unlisted ASA and thus remained regulated by the quota law.

[26]*The Economist* got this evidence wrong: "Companies fled the [Norwegian] stock market as quotas were faced in" (November 15, 2014).

next year ($t + 1$) and zero otherwise. Firms that convert are eliminated from the sample in the year of conversion because they are no longer ASA and hence have no option to convert. We include year fixed effects due to the strong downward time trend in $Shortfall$ as ASA comply with the quota regulation (Panel B of Figure 3). Our unbalanced panel contains 880 firm-years for 264 unlisted non-financial ASA, of which 150 convert over the period 2002–2009. While not tabulated, the coefficient estimate of γ_1 is a statistically insignificant 0.91 (full regression results available upon request). Thus, there is no evidence of a significant within-year (cross-sectional) correlation between the decision of an unlisted ASA to convert to AS and the firm's board gender composition.[27]

7 Conclusion

Several facts add power to an event study of the stock market reaction to Norway's board gender quota law: (1) As the first of its kind and initially resisted by the ruling political party, the quota law was highly unanticipated. (2) By covering all public limited companies (listed as well as unlisted ASA), it captures the bulk of the economy in market-value terms. (3) By imposing a severe penalty for noncompliance (forced liquidation), there was no uncertainty about universal compliance. (4) By regulating gender-balancing *only* (no other governance aspects), and because it was the result of a political decision unrelated to firm performance, inferences as to the causal effects of the law are relatively straightforward.

We revisit an important narrative emerging from prior research on this quasi-experiment: that the quota law caused an economically large decline in the market value of listed firms. We show that this narrative does not survive a closer scrutiny of the data. We reach this important conclusion not only within our larger sample and econometric analysis, but also after addressing the most relevant prior studies on their own terms. The narrative emerging from the evidence of this paper is that the quota constraint imposed statistically and economically negligible costs on regulated firms.

Our econometric adjustments of AD's event study result in two changes to their conclusion of a significantly negative market reaction: First, using AD's own estimation methodology, we show that AD's negative market reaction is in response to a second news event within their five-day event window that *decreased* (not increased) the likelihood of a quota law. This discovery effectively reverses AD's

[27]Bøhren and Staubo (2014) find that firms with a low fraction of female directors are more likely to convert *at some point* during the 2000–2009 period. However, we have verified that this correlation becomes insignificant once we include year fixed effects that control for the trend in the fraction of female directors caused by quota compliance.

own conclusion. Second, when we account for the contemporaneous cross-correlation of stock returns, which arises whenever the news event affects all firms simultaneously in calendar time, one can no longer reject the hypothesis of a zero market reaction over AD's five-day event window. We end our event-study analysis by providing new and comprehensive estimates of several quota-related news events and a cross-sectional analysis of individual (firm-level) abnormal returns. This analysis leads to one conclusion: that one cannot reject the null hypothesis of a zero stock market reaction to the quota law.

We also perform additional tests that corroborate the conclusion of a value-neutral effect of the quota. These include a long-short portfolio performance estimation (short in firms with pre-quota all-male boards), IV-analysis of post-quota changes in Tobin's Q and operating profitability (ROA), and legal conversions. We show that post-quota changes in Q and in ROA (the latter compared to large AS) are statistically indistinguishable from zero, except in the financial crisis year of 2008—changes that are most likely unrelated to the quota itself. Moreover, contrary to claims in the prior literature, over the total sample period, no listed ASA left the stock exchange for reasons other than M&A or bankruptcy and the number of listed firms actually increased during the implementation of the quota. Also, we show that the conversions undertaken by non-listed ASA are uncorrelated with the shortfall of female directors in the cross-section.

Overall, the evidence in this paper strongly suggests that firms and investors view Norway's forced gender-balancing law as a value-neutral regulatory constraint. A consistent explanation is that the supply of qualified female directors is sufficiently large to avoid a decline in firm value. Perhaps anticipating similar low-cost effects, several other western European countries have since decided to follow suit and adopted their own versions of board gender quotas.

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

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

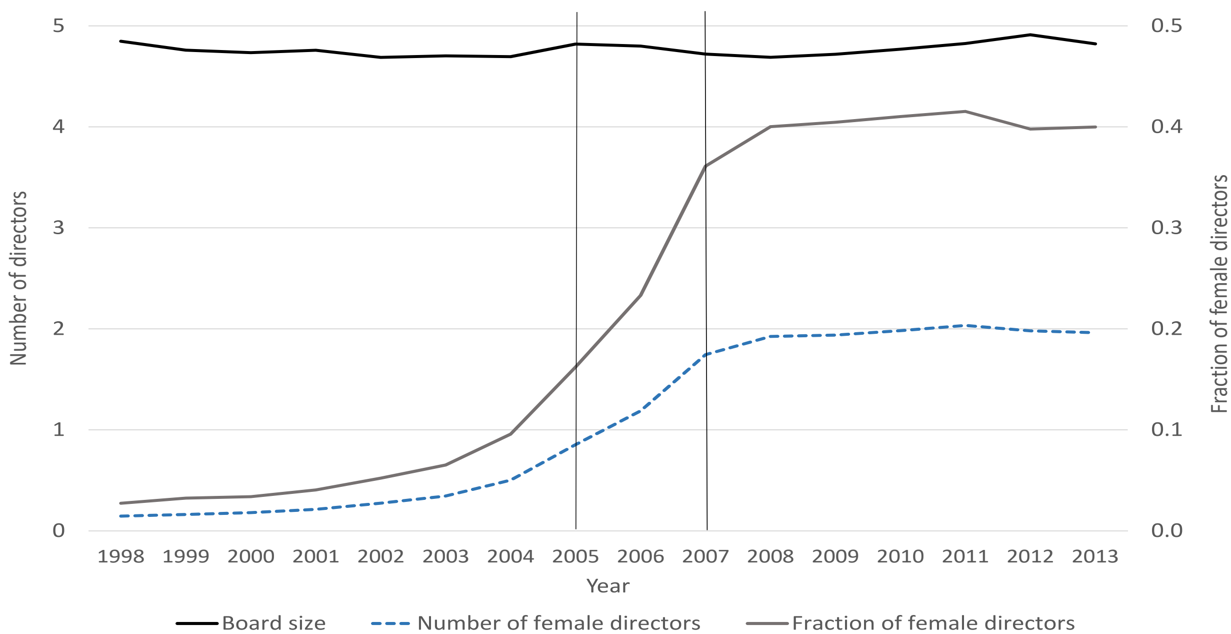
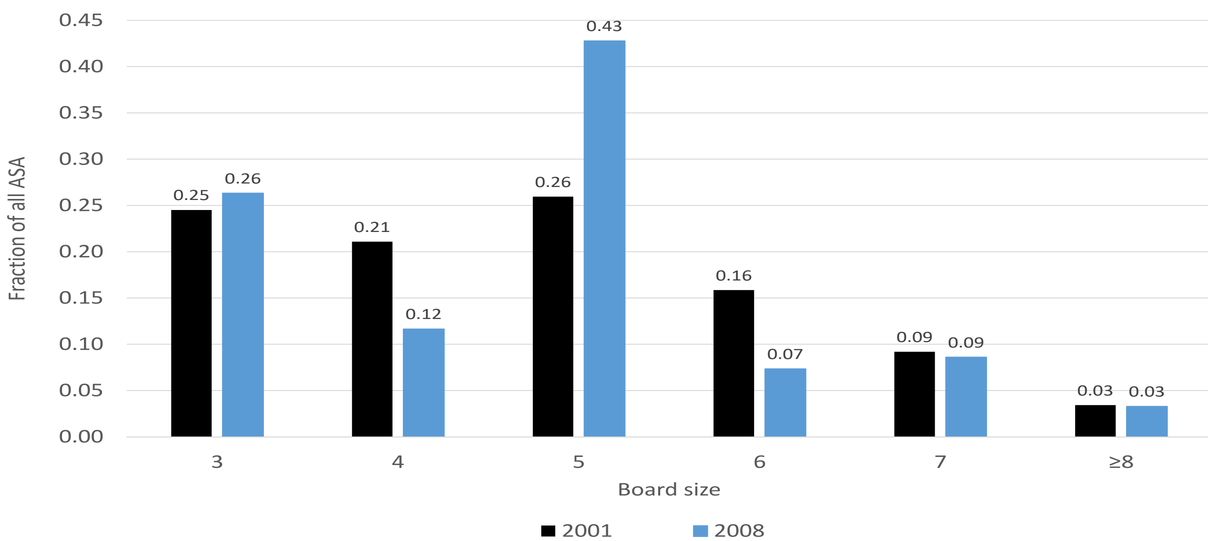


Figure 2: Frequency distribution of ASA board size and director board seats, 2001 and 2008

Panel A shows the frequency distribution of board size (the number of shareholder-elected directors). Panel B plots the frequency distribution of the total number of board seats in ASA and 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 (919 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 top 1% largest AS

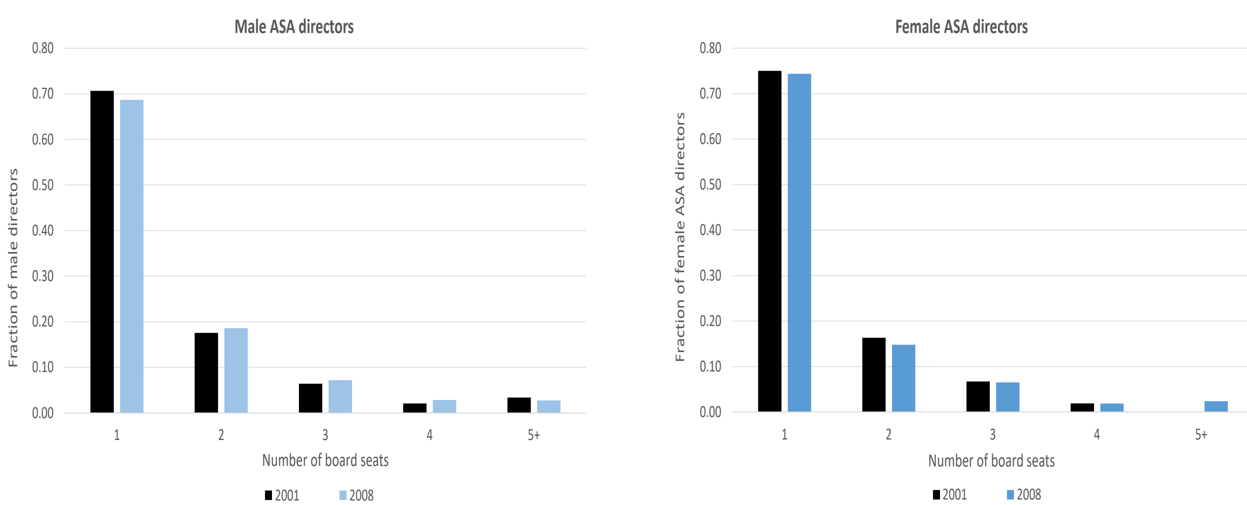
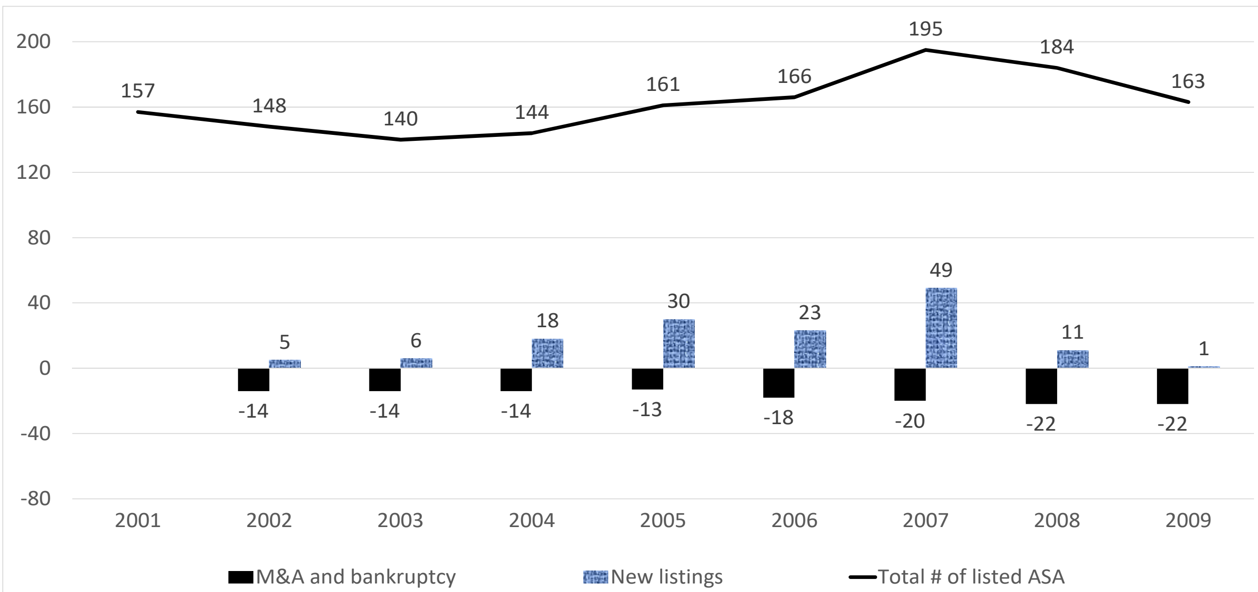


Figure 3: Total number and exits and entries of listed and unlisted ASA by year

The figure shows the total number of listed ASA (Panel A) and unlisted ASA (Panel B) at year-end, 2001-2009, and the number of exits and entries during the period 2002–2009. Firms enter and exit the legal form ASA by changing their bylaws, but typically give no reason for the change. No listed firm delist for reasons other than M&A or bankruptcy. Unlisted ASA exit because they are acquired or file for bankruptcy (192 firms), go public (51 firms), or for other reasons (156 firms). The sample is 288 listed and 467 unlisted non-financial ASA, 2002-2009. The data is from *Brønnøysund Register Centre*, complemented with manual searches of press releases and news for acquisitions by foreign firms.

Panel A: Listed ASA (no firm delists for reasons other than M&A and bankruptcy)



Panel B: Unlisted ASA (156 firms convert to AS for other reasons)

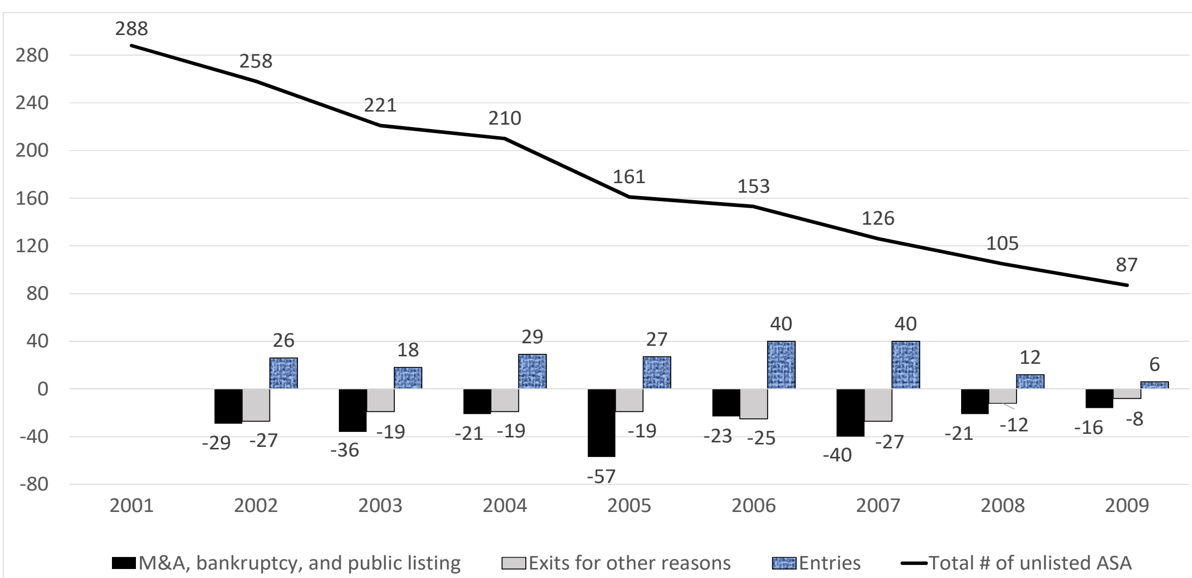


Table 1: NEWS EVENTS INCREASING THE PROBABILITY OF A BOARD GENDER QUOTA

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

Saturday, February 23, 2002: Gabrielsen 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. Importantly, the law proposal will be enacted with a sunset provision in 2005, canceling the quota if firms comply voluntarily by then. Cabinet promises compliance by government-owned firms within one year (*Dagens Næringsliv*).

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

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 and says that the sanction for non-compliance will likely be economic fines (*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 2: FEMALE DIRECTORS REQUIRED BY NORWAY'S BOARD GENDER QUOTA

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

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

Table 3: Firm and board characteristics for listed and unlisted ASA, 1998-2013

The table reports firm and board characteristics for listed (Panel A) and unlisted ASA (Panel B), 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 the board's directors with CEO experience fom an ASA or the largest 1% AS by revenue (about 1,000 firms per year), over the past three years. See Table 4 for definitions and data sources. The last row in each panel lists the pooled average across all firm-years, with the exception of the number of firms, which lists the average annual N over the sample period. The sample is 409 unique listed ASA and 888 unique unlisted ASA, of which 147 firms are listed in some years and unlisted in others.

Year	Number of firms (1)	Mean revenue (2)	Mean total assets (3)	% ownership of largest shareholder (4)	% female directors (5)	% female chairs (6)	% female CEOs (7)	Board CEO experience (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
2003	151	391	814	30	9.9	2.1	2.1	17.8
2005	174	446	910	31	22.2	1.7	1.1	17.0
2007	205	495	1098	33	38.9	2.9	2.9	18.5
2009	172	549	1127	36	41.1	4.1	3.5	17.7
2011	172	663	1236	34	41.5	10.5	4.1	17.2
2013	150	699	1373	34	41.6	8.8	4.7	15.6
Mean	174	501	1008	35	24.3	3.6	2.8	17.5
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
2003	346	74	226	56	5.0	2.8	5.7	16.2
2005	275	103	308	59	12.5	2.9	4.0	14.6
2007	260	93	369	58	33.9	6.9	6.2	13.4
2009	167	111	586	57	39.8	10.2	9.8	12.7
2011	126	197	651	55	41.5	11.1	13.6	15.0
2013	86	229	787	52	37.1	15.1	9.6	11.8
Mean	255	90	290	56	15.2	4.2	5.3	15.3

Table 4: DEFINITION OF VARIABLES USED IN THE ANALYSIS

The main data source is *Brønnøysund Register Centre*. Data for *Risk* and *Q* are from *Oslo Børsinformasjon*. Ownership is complemented with data from the Norwegian tax authorities (2004–2013). 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. 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.
ASA	Public limited company ("*Allmenaksjeselskap*"), regulated by the quota.
AS	Private limited company ("*Aksjeselskap*"), 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.
Board CEO experience	Fraction of directors with CEO experience from any ASA or one of the 1% largest AS by revenue, over the past three years.
Board busyness	The fraction of the board's directors that hold at least 3 board seats in ASA or the 1% largest AS by revenue.
Shortfall	Difference between the fraction of female directors required by the quota and that of the current board.
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 has zero female directors in 2001.
Pos_{2001}	Dummy equal to one if the firm has at least one female director in 2001, $Pos_{2001} = 1 - Zero_{2001}$.
Comply	Dummy equal to one in years $t \geq 2008$ (reflecting quota compliance by 12/2007).

Table 5: REPLICATING AND EXPANDING AD'S EVENT STUDY AROUND FEBRUARY 22, 2002

Panel A lists the average five-day $CAR(-2,2)$ around February 22, 2002 (our event 1 in Table 1) from AD's Table III. For each ASA, they estimate:

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

where r_i is the return of firm i (from Compustat Global), r_{imatch} is the average return to US-listed companies in firm i's GICS industry (from CRSP), and τ is the five days in the event window. Panels B1, B2, and C show our estimates of AD's average five-day $CAR_i(-2,2)$ using AD's sample, methodology, and (when possible) return data. Like AD, the p-values (in square brackets) are based on a standard error for the mean which assumes that all return observations are independent, $\sigma_N(\rho = 0) = \sigma/\sqrt{N}$. Panels B1 and B2 allows for multi-day returns, while Panels C and D require one-day returns in the event window, which reduces the sample size to 79 firms (only a subset of firms have a one-day return in Compustat Global on a given day in AD's event window). Except for Panel B2, the sample is split by firms' 2001 board-gender composition into $Zero_{2001}$ (column 2) and Pos_{2001} (column 3). Panel B2 replaces the firms in $Zero_{2001}$ (Pos_{2001}) with the smallest (largest) firms by revenue. Panel D reports the daily average $AR_i = (r_i - r_{imatch})_\tau$ on each day τ in the five-day event window. AD's board data is from annual reports, while our board data is from *Brønnøysund Register Centre*, where 69 firms (vs. 68 in AD) have zero female directors in 2001 ($Zero_{2001}$). The p-value in Column 4 uses a t-test. Significance levels: *** 1%, ** 5% and * 10%.

		All firms (1)	Firms with $Zero_{2001}$ (2)	Firms with Pos_{2001} (3)	Diff. in mean $Zero - Pos$ (4)
A: AD's own five-day average CAR (from AD's Table III)					
	$CAR(-2,2)$	-2.573***	-3.547***	-0.024	-3.523***
	p-value	[0.001]	[0.001]	[0.977]	[0.008]
	N	94	68	26	
B1: Estmating AD's five-day CAR allowing multi-day returns					
	$CAR(-2,2)$	-2.827***	-3.894***	0.115	-4.009***
	p-value	[0.001]	[0.000]	[0.913]	[0.007]
	N	94	69	25	
B2: Replacing $Zero_{2002}$ (Pos_{2002}) with the smallest (largest) firms					
	$CAR(-2,2)$	-2.827***	-3.658***	-0.533	-3.125*
	p-value	[0.001]	[0.000]	[0.724]	[0.082]
	N	94	69	25	
C: Estimating of AD's five-day CAR requiring one-day returns					
	$CAR(-2,2)$	-2.320**	-3.257***	0.106	-3.363*
	p-value	[0.017]	[0.010]	[0.930]	[0.054]
	N	79	57	22	
D: Daily AR estimates w/AD's methodology and one-day returns					
Feb 20 (Wed)	$AR(-2)$	-0.939*	-1.154*	-0.345	-0.810
	p-value	[0.053]	[0.060]	[0.634]	[0.388]
	N	64	47	17	
Feb 21 (Thu)	$AR(-1)$	0.416	0.037	1.626**	-1.589*
	p-value	[0.417]	[0.955]	[0.013]	[0.070]
	N	67	51	16	
Feb 22 (Fri)	$AR(0)$	-0.753	-0.695	-0.887*	0.192
	p-value	[0.254]	[0.453]	[0.099]	[0.855]
	N	66	46	20	
Feb 23 (Sat): Reversal announcement. OSE closed, no trading.					
Feb 25 (Mon)	$AR(1)$	-1.995***	-2.537***	-0.711	-1.826**
	p-value	[0.000]	[0.000]	[0.156]	[0.026]
	N	64	45	19	
Feb 26 (Tue)	$AR(2)$	0.405	0.286	0.672	-0.385
	p-value	[0.402]	[0.657]	[0.291]	[0.667]
	N	65	45	20	

Table 6: ADJUSTING AD'S EVENT-STUDY FOR RETURN CROSS-CORRELATION

The table revisits AD's event study by using their sample and data sources, but adjusting for the cross-correlation of returns through a portfolio estimation of CAR(-2,2) around February 22, 2002 (event 1 in Table 1). We estimate a time-series regression of the model:

$$r_{pt}^{-I} = \alpha_p + AR_p d_t + \epsilon_{pt}.$$

The dependent variable $r_{pt}^{-I} \equiv \frac{1}{N}\sum_{i=1}^{N}(r_i - r_{imatch})_t$ is the equal-weighted portfolio of industry-matched returns, where r_i is the return of ASA i (from Compustat Global) and r_{imatch} is the average return to US-listed companies in firm i's GICS industry (from CRSP) on day t, and d_t is a dummy that takes the value of one for all days in the five-day event window (-2,2). Panel A shows the coefficient estimates $CAR(-2,2) = 5AR_p$ from the time-series regression. The p-values (in square bracket) use the standard error from the regression, $\sigma_{5AR_p}(\rho) = 5\sigma_{AR_p}(\rho)$, which accounts for the cross-correlation in returns. Panel B report the daily AR in the five-day event window, using the same portfolio estimation but with five different dummies d_t (one for each day). We require a firm to have at least a single one-day return in the event window. This reduces the sample size to 79 firms in Panel A (from AD's 94). The number of observations is lower in Panel B because only a subset of firms have a one-day return in Compustat Global on a given day in the event window. The board data is from *Brønnøysund Register Centre*. The sample is split by firms' 2001 board-gender composition into zero female directors ($Zero_{2001}$, column 2) and at least one female director (Pos_{2001}, column 3). Significance levels: *** 1%, ** 5% and * 10%.

		All firms (1)	Firms with $Zero_{2001}$ (2)	Firms with Pos_{2001} (3)	Diff. in mean $Zero - Pos$ (4)
A: Time-series portfolio estimation of the five-day CAR (one-day returns)					
	$CAR(-2,2)$	-2.203	-3.364	0.924	-4.288
	p-value	[0.521]	[0.356]	[0.796]	[0.116]
	N	79	57	22	
B: Time-series portfolio estimation of the daily AR (one-day returns)					
Feb 20 (Wed)	$AR(-2)$	-0.806	-1.014	-0.231	-0.784
	p-value	[0.598]	[0.532]	[0.885]	[0.518]
	N	64	47	17	
Feb 21 (Thu)	$AR(-1)$	0.549	0.176	1.740	-1.563
	p-value	[0.720]	[0.913]	[0.276]	[0.197]
	N	67	51	16	
Feb 22 (Fri)	$AR(0)$	-0.620	-0.555	-0.773	0.218
	p-value	[0.685]	[0.732]	[0.628]	[0.857]
	N	66	46	20	
Feb 23 (Sat): Reversal announcement. OSE closed, no trading.					
Feb 25 (Mon)	$AR(1)$	-1.862	-2.397	-0.597	-1.800
	p-value	[0.224]	[0.140]	[0.708]	[0.138]
	N	64	45	19	
Feb 26 (Tue)	$AR(2)$	0.538	0.426	0.785	-0.359
	p-value	[0.725]	[0.793]	[0.622]	[0.767]
	N	65	45	20	

Table 7: Abnormal returns to portfolios of OSE-listed firms on key quota event dates

The table reports cumulative abnormal stock returns, $CAR(-1,0) = 2AR_p$, for portfolios of OSE-listed firms, estimated using the return-generating process:

$$r^e_{pt} = \alpha_p + \beta_p r^e_{wt} + AR_p d_t + \varepsilon_{pt},$$

where r^e_{pt} 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_t is a dummy for the event window (-1,0), and r^e_{wt} is the daily excess return on the MSCI stock market world index. The five events are defined in Table 1. High shortfall firms have a female director $Shortfall$ (fraction of additional female directors required by quota, see Table 1) at or above the median in the preceding year. 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 one-day return observations in both days in the event window and ≥100 observations in the estimation period (-252, -2). We exclude any earlier event date in the estimation period. For $CAR_{1-5}(-1,0) = 10AR_{1-5}$, 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. 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							
$CAR(-1,0)$	-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							
$CAR(-1,0)$	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							
$CAR(-1,0)$	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							
$CAR(-1,0)$	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							
$CAR(-1,0)$	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)							
$CAR_{1-5}(-1,0)$	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 8: 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, $k = 1, \ldots, 5$ (Table 1). For each firm i, the daily average abnormal return AR_{ik} is estimated for each event k using the regression model in Table 7. 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*, *Largest owner*, *Government control*, *Codetermination*, *Risk*, and *Total assets*, and κ_i is industry fixed effects. $Shortfall$ is the fraction of female directors missing to fill the quota requirement. All variables are from the year-end prior to the event and defined in Table 4. A constant is included, but not reported. The sample is OSE-listed ASA. Daily stock returns are from *Oslo Børsinformasjon*. Firm and board characteristics are from *Brønnøysund Register Centre*. Robust standard errors (White estimator) are reported in parenthesis. Stars indicate significance levels: *** 1%, ** 5%, * 10%.

	Date of quota news event				
	22-Feb-2002 (1)	8-Mar-2002 (2)	16-Jun-2003 (3)	1-Dec-2005 (4)	9-Dec-2005 (5)
$Shortfall$	-0.051	0.017	0.001	-0.013	-0.002
	(0.067)	(0.055)	(0.061)	(0.022)	(0.023)
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	0.125	0.157	0.152	0.108	0.134
N (firms)	129	131	123	132	133

Table 9: ABNORMAL PERFORMANCE OF PORTFOLIOS CLASSIFIED BY FEMALE REPRESENTATION

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 100 and 32, respectively. In columns (1) - (3), the abnormal stock return is estimated using the following three factor return-generating process:

$$r^e_{pt} = \alpha_p + \beta_{p1} r^e_{wt} + \beta_{p2} HML_t + \beta_{p3} SMB_t + \varepsilon_{pt}, \quad t = 2/2002, ..., 4/2008,$$

where r^e_{pt} is the monthly USD-denominated stock return to portfolio p of domestic OSE-listed ASA, converted to USD using the monthly exchange rate, in excess of the current month's 3-month US Treasury bill. r^e_{Wt} 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_p	-0.002	0.003	-0.004	-0.001	0.004	-0.005
	(0.006)	(0.005)	(0.004)	(0.006)	(0.005)	(0.004)
W^e	1.418***	1.416***	0.002	1.406***	1.373***	0.034
	(0.154)	(0.121)	(0.109)	(0.164)	(0.128)	(0.116)
HML	-0.134	0.320	-0.454	-0.120	0.373	-0.493
	(0.426)	(0.336)	(0.302)	(0.434)	(0.340)	(0.306)
SMB	1.119***	0.721***	0.398*	1.139***	0.794***	0.345
	(0.309)	(0.244)	(0.219)	(0.324)	(0.254)	(0.229)
MOM				-0.041	-0.151	0.111
				(0.188)	(0.148)	(0.133)
R^2	0.602	0.683	0.060	0.602	0.687	0.069
Observations (months)	75	75	75	75	75	75

Table 10: IV REGRESSIONS FOR TOBIN'S Q

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 firm and year fixed effects, respectively. $\widehat{Shortfall}$ is the fitted value from the first-stage IV regressions reported in the odd-numbered columns in Panel B:

$$Shortfall_{it} = \alpha + Base_i \sum_{t=2004}^{2009} \beta_t \tau_t + \theta_i + \tau_t + u_{it},$$

The even-numbered columns in Panel B report the reduced form regression, with Q_{it} as the dependent variable. The variable $Base$ is the 2002 shortfall of female directors, $Shortfall_{2002}$ (columns 1, 2, 5 and 6), log of Sales in 2002 (columns 3 and 4), and the 2001 shortfall of female directors, $Shortfall_{2001}$ (columns 7 and 8). The sample is 167 unique OSE-listed ASA, 2003-2009, and 820 firm-years. Five government-controlled firms are excluded in columns (5) and (6). Standard errors clustered by firm are reported in parenthesis. Significance levels: *** 1%, ** 5%, and * 10%.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Sample:	All firms		All firms		Ex. gov-contr. firms		All firms	
A: Second-stage IV regression for Q_{it}								
$\widehat{Shortfall}$	1.912**		6.105**		1.680		0.694	
	(0.833)		(2.495)		(1.053)		(1.236)	
F	15.90		12.01		15.66		16.76	
R^2	0.128		-0.197		0.136		0.141	
N	820		734		785		790	
B: First-stage IV regression for $Shortfall_{it}$ and reduced form regression for Q_{it}								
Base:	$Shortfall_{2002}$		$log(Sales)_{2002}$		$Shortfall_{2002}$		$Shortfall_{2001}$	
Regression:	First-stage IV	Reduced form	First-stage IV	Reduced form	First-stage IV	Reduced form	First-stage IV	Reduced form
Dep. variable:	$Shortfall_{it}$	Q_{it}	$Shortfall_{it}$	Q_{it}	$Shortfall_{it}$	Q_{it}		
$Base * \tau_{2004}$	-0.121	0.569	-0.005	-0.039	-0.106	0.608	-0.080	0.604
	(0.075)	(0.362)	(0.004)	(0.025)	(0.099)	(0.452)	(0.087)	(0.448)
$Base * \tau_{2005}$	-0.420***	1.025	-0.010*	-0.127	-0.478***	0.974	-0.274**	-0.427
	(0.103)	(0.873)	(0.006)	(0.085)	(0.129)	(1.134)	(0.114)	(0.960)
$Base * \tau_{2006}$	-0.531***	0.244	-0.006	0.099	-0.560***	0.356	-0.450***	-0.386
	(0.101)	(0.743)	(0.008)	(0.060)	(0.135)	(1.010)	(0.134)	(0.630)
$Base * \tau_{2007}$	-0.796***	-0.823	0.023***	0.136**	-0.791***	-0.750	-0.546***	-0.079
	(0.080)	(0.672)	(0.007)	(0.066)	(0.117)	(0.945)	(0.133)	(0.698)
$Base * \tau_{2008}$	-0.906***	-1.591**	0.026***	0.183***	-0.957***	-1.539	-0.627***	-0.340
	(0.086)	(0.669)	(0.007)	(0.068)	(0.117)	(0.954)	(0.132)	(0.835)
$Base * \tau_{2009}$	-0.964***	-1.149	0.028***	0.172**	-1.049***	-1.112	-0.634***	0.065
	(0.091)	(0.698)	(0.008)	(0.078)	(0.120)	(0.994)	(0.131)	(0.899)
F	85.88	10.97	44.08	12.05	87.05	10.85	45.54	10.93
R^2	0.660	0.164	0.640	0.208	0.665	0.164	0.624	0.144
N	829	829	742	748	794	794	799	799

Table 11: QUOTA-INDUCED CHANGES IN ROA

Columns (1)–(4) report 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 $t \geq 2008$ and zero otherwise. The vector $\mathbf{X}_{it}$ contains the following firm characteristics: *Firm age*, *Size*, *Leverage*, *Largest owner*, *Board size*, *Board CEO experience*, *Board busyness*, and a constant (all suppressed). The variables are defined in Table 4. In columns (5) and (6), the model is:

$$ROA_{it} = \gamma_0 + ASA_i \sum_{2008}^{2013} \gamma_t \tau_t + \gamma_2 \mathbf{X_{it}} + \theta_i + \tau_t + \epsilon_{it},$$

The sample comprises 436 ASA (treated firms) and 1,786 of the 1% largest AS by revenue (control firms), 2003-2013. We exclude financial firms and ASA registered as AS at some point during the sample period, and require firms to have at least two observations and no missing values for the control variables. The sample period is 2003–2009 in columns (1)–(2) and 2003–2013 in columns (3)–(6). Standard errors clustered by firm are reported in parenthesis. Stars indicate significance levels: *** 1%, ** 5%, and * 10%.

	2003-2009		2003-2013			
	(1)	(2)	(3)	(4)	(5)	(6)
$ASA * Comply$	-0.024**	-0.022**	-0.012	-0.000		
	(0.012)	(0.011)	(0.011)	(0.010)		
$ASA * \tau_{2008}$					-0.049***	-0.042***
					(0.016)	(0.015)
$ASA * \tau_{2009}$					-0.015	-0.011
					(0.015)	(0.014)
$ASA * \tau_{2010}$					0.002	0.018
					(0.017)	(0.016)
$ASA * \tau_{2011}$					-0.009	0.008
					(0.021)	(0.019)
$ASA * \tau_{2012}$					0.004	0.023
					(0.020)	(0.017)
$ASA * \tau_{2013}$					0.025	0.039**
					(0.017)	(0.017)
Firm characteristics	No	Yes	No	Yes	No	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.024	0.135	0.018	0.122	0.022	0.127
N (firm-years)	7241	7241	11,354	11,354	11,354	11,354

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