Insider trading, risk aversion, and gender[*]

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Abstract

We analyze three decades of primary insiders' non-routine trades on the Oslo Stock Exchange and test for gender-based differences in risk-aversion and access to inside information. Long-run returns-based and holdings-based performance measures are statistically insignificant. However, there is some evidence of a positive female-specific director network information effect of Norway's board gender-balancing law, which significantly increased female director networks. Female insider purchases increased significantly immediately after the financial crisis, both absolutely and in relative terms, suggesting that these female directors and executives are no more risk-averse than their male colleagues.

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"If women must be more like men to break the glass ceiling, we might expect gender differences to disappear among directors."
— Renée Adams and Patricia Funk, *Management Science*, 2012 (abstract).

1 Introduction

As surveyed by Croson and Gneezy (2009), the literature on gender differences suggests that there are systematic dispositional differences between males and females. For example, data from laboratory settings, where the participants are typically students or workers, tend to indicate that females are more risk-averse than males (Eckel and Grossman, 2008; Sapienza, Zingales, and Maestripieri, 2009). On the other hand, Adams and Funk (2012) argue that this type of evidence may not carry over to the more select group of individuals in corporate leadership, who are rarely available for such experiments. That is, to be considered a candidate for a board seat in a male-dominated public corporation, females may have to develop core values and risk attitudes that are similar to male directors (above quote). After surveying directors in Swedish listed companies in year 2005, Adams and Funk conclude that female executives and directors are, if anything, somewhat *less* risk averse than their male counterparts.

We analyze gender-based differences in primary insiders' risk aversion and access to inside information as evidenced more directly by the level and performance of their non-routine insider trades. Notwithstanding insider-trading law enforcement, primary corporate insiders will from time to time possess price-sensitive information that is not yet publicly available and that may form the basis for a decision to trade. The central question that we address is whether the level and change in insider holdings, trading propensities and performance differ significantly across male and female insiders, and whether such differences (if any) may be traced back to differences in risk aversion and access to information. Essentially, we move the analysis of gender-based differences in risk aversion from the traditional survey data and into the stock-market arena. Risk aversion undoubtedly plays a role in primary insiders' investment decisions with the potential to significantly affect both wealth and reputation.

Our empirical analysis covers three decades of population data on primary insider stock holdings and trades on the Oslo Stock Exchange (OSE). A study of insider performance on this exchange is of particular interest because (i) insiders own a relatively large percentage of the total number of outstanding shares, and (ii) the exchange is dominated by relatively small, high-volatility companies. Both (i) and (ii) are factors that may contribute to profitable trades based on inside information. Moreover, our population data allows us to form a unique portfolio using actual insider holdings—labelled the 'insider portfolio'. While not a managed portfolio (it aggregates individual insiders' trading decisions), if primary insiders as a group tend to 'buy low and sell high' based on inside information, then this portfolio should exhibit positive abnormal performance.

Our analysis makes contributions in three related areas. We begin with a battery of tests for gender-based differences in trading propensity and abnormal insider performance across our total sample period, 1986–2016. This performance analysis implements several modern econometric techniques, ranging from the returns-based approach that is common in the literature on mutual fund performance (Ferson, 2010; Wermers, 2011) to the much less explored holdings-based approach, which requires data on individual investors' stock holdings (Grinblatt and Titman, 1993; Eckbo and Smith, 1998; Ferson and Khang, 2002). We also include the more traditional estimates of the information content of insider trades in the form of the short-term market reaction, which began with Jaffe (1974) and Seyhun (1986). The combination of three decades of population data on insider holdings and modern approaches to performance measurement makes this an unusually comprehensive study of insider trades.

It has been argued that informal male networks may disadvantage females, especially in organizations where they are under-represented (Moore, 1988; Lyness and Thompson, 2000; Inci, Narayanan, and Seyhun, 2017). Our second area of contribution is to test this argument using the exogenous shock to the nation-wide primary-insider network caused by Norway's board gender-balancing law. In December of 2005, this law required all domestic public limited liability companies (listed as well as unlisted) to increase the percent female directors to 40% over the two-year period 2006–2007 (or else be liquidated!).[1] All firms were in full compliance by early 2008, and we test whether this dramatic expansion of the number of female primary insiders impacted the holdings and trading performance of female directors.

Our third and final area of contribution involves using the exogenous shock to market values during the 2008 financial crisis to test for gender-based differences in primary insider trading propensities. We argue that the market crash had two potential effects on the incentives for insiders to trade, both of which depend on the individual's risk aversion. The first incentive is to rebalance the insider's overall investment portfolio so as to restore the optimal allocation between investments in risky and risk-free assets. The second incentive is to bet against the stock market (purchase more shares) if the insider believes that

[1]The percent female directors in 2005 was 15% on average. See Ahern and Dittmar (2012), Eckbo, Nygaard, and Thorburn (2019) and Bertrand, Black, Jensen, and Lleras-Muney (2019) for discussions of the quota-induced board changes and their economic effects.

the market crash has caused the firm's stock to become undervalued.[2] Under the conventional view that females are generally more risk averse than males, female insiders should trade less than male insiders after the crash, which is what we test.

Our main empirical conclusion can be briefly summarized as follows: First, without explicitly conditioning on Norway's mandatory board gender-balancing or on the financial crisis, abnormal performance is statistically insignificant across the entire battery of tests, as are gender-based performance differences. Second, we find that the gender-quota law causes the short-term market reaction to female insider purchases to become positive and statistically significant—and similar to that of male insider purchases. No such improvement in the market reaction is observed for male insiders whose network declined with the influx of female directors. This finding suggests that the substantial increase in the female director network has increased the amount of inside information available to female directors, which in turn causes the market to react more strongly to female directors' trades.

Third, under the conventional view that females are generally more risk averse than males, female insiders should trade less than male insiders after the crash. However, we find the opposite. Conditional on a number of firm-specific characteristics that may also affect insider trades (such as trading costs and stock volatility), we show that female primary insiders not only increase their purchases in the stocks where they are insiders, they are also more likely than male primary insiders to buy shares. This evidence contradicts the notion that females are more risk averse than males—as far as primary insiders go. As such, our evidence also supports the argument of Adams and Funk (2012) that the selection of females as executives and directors may require these individuals to be "more like men to break the glass ceiling" (above quote).

In terms of the literature, Inci, Narayanan, and Seyhun (2017) is to our knowledge the only other study examining gender-based information issues in the context of insider trading. They use US data where the fraction of female core insiders remains relatively low throughout their sample period (1975–2012). They provide standard estimates of the cumulative market reaction to insider trades, which is not unlike our Section 6 where we present estimates of abnormal stock returns around trading dates. While Inci, Narayanan, and Seyhun (2017) do not examine network shocks, their evidence of a positive short-term market reaction to both female and male primary insider purchases is interesting. Controlling

[2]For discussions of how experiencing market crashes may affect trading, see e.g. Malmendier and Nagel (2011), Weber, Weber, and Nosić (2012), Hoffmann, Post, and Pennings (2013), and Guiso, Sapienza, and Zingales (2018).

for the insiders' formal status in the firm, their estimates suggest that female executives likely have an informational disadvantage relative to males. Notwithstanding differences in data, institutional setting and empirical methods across the two studies, this suggestion is certainly consistent with our own finding that the positive shock to Norway's female director network likely increased female directors' access to inside information.

The rest of the paper is organized as follows. Section 2 describes our insider trading data and provide trade summary statistics. Section 3 compares returns-based stock market performance of portfolios formed from trades of female and male insiders, respectively, while Section 4 presents insider performance using holdings-based metrics. In Section 5, we implement an event study to identify the short-term market reaction to insider trades, again classified by gender. In Section 6, we turn to the potential effect of the board reform on insider performance, while Section 7 provides evidence on gender-based trading activity during the financial crisis. Section 8 concludes the paper.

2 Regulations, data and sample characteristics

2.1 Insider trading regulations

Our sample period covers the three decades 1/1986–12/2016, where the first decade were ruled by Norway's first-generation insider trading laws. While insiders were not required by this law to report their trades, the OSE demanded such reporting on a monthly basis and made the reports publicly available. In 1996, Norway passed new legislation ("Lov om Verdipapirhandel") adopting the European Union's (EU's) principles for insider trading regulations.[3] The law defines *primary* insiders as top management (including the chief executive officer and chief financial officer) and board members. Our main empirical analysis is based on these primary insiders. Under the 1996 legislation, insiders publicly report their trades within one day, which typically happens prior to next day's opening. The law also specifies certain insider trade blackout periods, including prior to corporate earnings announcements.

As is typical in the literature on insider trades, we study *all* trades based on public reporting—not just the trades that were judged to be illegal *ex post* (Meulbroek, 1992; Bhattacharya, 2014). What

[3]Norway is under treaty obligation to adopts EU regulations, including EU restrictions on insider trades. There has been largely minor adjustments to EU's and Norways insider trading regulations between 1997 and the end of our sample period in 2016. For a summary of Norway's insider trading regulations, see sections 6 (Issuer's obligations) and 7 (primary insider's obligations) of NOU 2017:4.

is clear *ex ante* is that insiders do from time to time possess price-sensitive information, and that the likelihood that they trade on this information depends on the expected financial and reputational cost of doing so. While insider trading law enforcement was relatively weak in the first decade of our sample period (Eckbo and Smith, 1998), it grew much stronger in our sample period after the adaptation of the EU regulation, which forms the core period of our analysis. A quick search identifies a total of 22 court cases where the defendant is charged with criminal insider trading over the period 1998–2018. In general, a conviction leads to both jail time and a fine equal to the estimated trading profit. This enforcement likely deters blatantly illegal trades, but leaves room for smaller information-based trades that are hard to classify as illegal *ex post*. The performance tests reported below examine whether insiders on average are able to include, undetected, such smaller information-based trades among their more routine trades.

The stricter enforcement that came with the adoption of EU rules may also have lowered the incentive to hold shares as an insider. There is some evidence of this in Figure 1, which shows a significant decline in the average quarterly percent inside ownership on the OSE after 1992. The figure is based on all (not just primary) reporting insiders and is computed by averaging, each quarter, the percent insider holdings across all OSE listed firms. A second important reason for the decline is the substantial rise in the OSE market value after 1992, as shown by the right-hand-side axis in Panel A of Figure 1. The post-1992 period also saw the OSE-listing of some large companies, including the partial privatization and OSE listing of government-controlled companies such as Telenor and Equinor (previously Statoil). Panel B of the figure shows that the average percent insider holding is typically higher for relatively small companies, which prompts us to form portfolios of insider holdings weighted by value invested as well as by shares owned (details in Section 3.1).

2.2 Data sources and sample characteristics

For trades prior to 1997, we rely on the monthly OSE reports, which include the name of the insider and the quantity bought or sold over the month.[4] Beginning in 1997, we collect insider trades from 1997 from OSE electronic records (`https://newsweb.oslobors.no/`). To be included, the trade announcement must contain the name and formal company-position of the insider, the trade date, and the number of shares bought or sold. The report typically also include the insider's share holding after the reported

[4]If the OSE report prior to 1997 omits the exact trading date, we use the date of the OSE report, which suffices for the monthly time-series estimation in sections 3 and 4 below. After 1996, the exact trading dates are always reported.

trade. As detailed in Section 3.1 below, if the holding is not reported, we reconstruct the holding by adding or subtracting the month's purchases or sales to the previous month's holding (before removing routine trades).

In addition to these insider trading reports, we obtain data on financial information such as stock prices, accounting information and corporate events from the OSE data service and Datastream, interest rates from Norges Bank (The Norwegian Central Bank), and other macroeconomic information from the Norwegian Bureau of Statistics (SSB). Moreover, we collect firm-level data on board size, board composition and director gender from the Brønnoysund Register Centre (1998–2014) via the Norwegian School of Economics (Berner, Mjøs, and Olving, 2013). Figure 2 shows the evolution of board size and the number and percentage of females on the boards of OSE-listed Norwegian firms over the 1998–2014 period. In 1998, the percentage females was 10%, which increased to 15% just prior to 2005 and to 40% over the following two years as firms complied with the board gender-quota law. This increase led to a dramatic expansion of the director network available to female directors, which is the instrument that we use to identify a possible female network-information effect in Section 6.

As shown in the first row in Panel A of Table 1, our data includes a total of 47,429 insider transactions over the 1986–2016 sample period. This total transaction count includes trades in different firms by the same insider. This, however, happens only rarely: Eckbo, Nygaard, and Thorburn (2019) show that the dispersion of board seats is very high both before and after the forced gender-balancing. For example, in 2008, 75% of all directors hold a single board seat only, with an additional 15% holding two seats only. While the total transactions is evenly split between the two sub-periods that are separated by the 1996 insider trading legislation, trades in the second subperiod 1997–2016 are the main focus of our empirical tests. We include performance estimates based on the earlier subperiod primarily for comparison purposes to the results in Eckbo and Smith (1998), who do not classify insiders by gender.

The second row in Panel A of Table 1 shows that we succeed in classifying 38,504 of the 41,429 transactions by gender.[5] As per the third row, 56% (21,663) of the gender-based insider trades are executed by primary insiders (management and board members), which increases from 33% in the early subperiod to as much as 75% after 1996.[6] The fourth row of Panel A shows that 88% (19,108) of the

[5]While not tabulated, this represents a classification success rate of 81% over the thirty-year sample period, with a success rate of 88% in the most recent sub-period. We identify the insider's gender from his or her given name, which in Norway nearly always identifies the gender. For insiders with foreign names, we include only those where the gender is unambiguous from the given name.

[6]This increase is in part a reflection of our attempt to backfill the insider's corporate position in the 1986–1996 data,

gender-based primary insider trades are non-routine. Here we filter out routine (repeat) trades using the method suggested by Cohen, Malloy, and Pomorski (2012). An insider trader in month t is labeled a "repeat performer" if the same insider traded in the same calendar month in each of the three years preceding the trade in month t. The elimination of repeat trades reduces the total number of trades by gender-based primary insiders by 12%, from 21,663 to 19,108. Interestingly, this reduction in sample size is substantially less than the 50% reduction reported by Cohen, Malloy, and Pomorski (2012) for US insider trades. The likely reason for this difference is the low frequency of stock- and option-based executive compensation plans among OSE companies.[7]

Next, panels B and C of Table 1 describe differences in trading activity between male and female insiders—primary as well as all insiders. In Panel B, our inside transactions covers 649 OSE-listed firms over the 1986–2016 sample period. Of these, 466 firms have reported trades by female insiders, while there are only three firms without any trades by male insiders. Of the of 16,473 distinct insiders, 18% (3,003) are female, while 17.2% of the 5,967 distinct primary insiders are female.

The transaction value over the sample period totals NOK 161.6 billion for purchases and 84.9 billion for sales, all measured in 2016 constant kroner.[8] Of these totals, female insiders undertake only 1.1% of the value of purchases and 2.4% of the value of sales. In terms of individual transactions, however, females undertake a much larger percentage: 13.2% of the total of 38,504 purchases and sales in Panel A are by female insiders. The distribution of NOK trade size is highly skewed. The value in NOK per transaction is smaller on average for female insiders than for males. For primary insiders, the median female NOK purchase size is about half that of the median male purchase.

In Panel C of Table 1, we follow Inci, Narayanan, and Seyhun (2017) and report, for each insider, the average annual number and value of his/her trades per year over the insiders' tenure period. This measure is not affected by the low fraction of female insiders early in our sample period, and therefore provides a more direct comparison of the trading intensities of male and female insiders. In this calculation, the first year of an insider's tenure period is the year of the first reported trade in our data, while the ending year is the year of the last reported trade. Thus, an insider with just one reported trade—or several

when the insider trading reports did not reveal this information. We perform this backfilling manually using post-96 position information as well as comprehensive directorship data. However, this identification is likely somewhat less comprehensive than is the case for the 1997–2016 period where we are able to rely on the insider trading reporting itself.

[7]Prior to 1999, stock options as a form of managerial compensation was extremely tax disadvantaged: the exercise value was taxed as regular income in the year of the option grant.

[8]Deflation uses the CPI from the Norwegian Bureau of Statistics (SSB).

trades within one year—are recorded as having a tenure period of just one year. The results in Panel C show that male insiders tend to trade more in total NOK. However, trading *intensity*—the number of transactions per year over the insider's tenure period—is similar across male and female insiders, whether the insider is primary or not.

Turning to trades by primary insiders, the main focus of our empirical analysis, Figure 3 shows the annual percent of all primary insider trades that are performed by female executives and directors. Female executives trade substantially more than female directors. Moreover, as expected, the percentage of all primary female trades jumps following the 2005 board quota law and relatively more so for female directors. Table 2 provides additional information on this trend in terms of purchases and sales and trade size.

We next present a battery of long-term performance tests using portfolios of primary insider holdings, as well as estimates of the short-term market reaction to primary insider trades. We first provide commonly used returns-based metrics (Section 3) before switching to the much less common holdings-based measures (Section 4). This is then followed by estimates of the short-term market reaction to individual insider trades—event-study type of analysis (Section 5). We subsequently use all three abnormal return metrics (returns-based, holdings-based and even-study based) to examine potential female insider network information effects on trading performance (Section 6).

3 Returns-based portfolio performance of primary insiders

Below, we present estimates of the monthly abnormal performance of primary insider portfolios in calendar time (1986–2016) using returns-based metrics that are common in the extant literature on insider trading and mutual fund performance. We first details the insider portfolio formation and then describe the returns-based performance metrics and evidence.

3.1 Insider portfolio formation

We construct two sets of portfolio weights using the $N = 649$ OSE-listed firms with reported primary insider trades over the period of $T = 372$ months, 1986-2016. Let N_t denote the total number of OSE-listed firms in month t and ω_{it}^k the weight of firm i in insider portfolio k at time t. The two sets of portfolio weights are defined as follows:

$$\omega_{it}^{k} = \begin{cases} \omega_{it}^{ow} \equiv (s_{it}/S_{it})/\sum_{i=1}^{N_t}(s_{it}/S_{it}) & \text{(insider-ownership-weights)} \\ \omega_{it}^{vw} \equiv h_{it}/\sum_{i=1}^{N_t} h_{it} & \text{(insider-value-weights)} \end{cases} \tag{1}$$

S_{it} denotes firm i's total number of shares outstanding in month t, of which insiders hold s_{it} shares. This insider holding is worth $h_{it} = p_{it}s_{it}$, where p_{it} is firm i's stock price at time t. Hence, ω_{it}^{ow} gives greater weights to firms in which insiders hold a larger ownership share of the outstanding stock, while ω_{it}^{vw} gives greater weight to firms with relatively large absolute value of insider investment.[9]

To construct ω_{it}^{ow} and ω_{it}^{vw}, we use the insider holdings (number of shares) contained in the insider reports to the OSE. If the holding is not reported, we reconstruct the holding by adding or recursively subtracting the month's purchases or sales to the previous month's holding. In this reconstruction, we adjust for changes in the number of shares outstanding caused by stock splits. Moreover, absent information to the contrary, we assume that insiders purchase their pro rata share of stock issues that use the rights-offering method, which is commonly used in Norway (Bøhren, Eckbo, and Michalsen, 1997). Also, if a firm with positive insider holdings delist from the stock exchange, we assume that the insider's holding is brought to zero (sold) at the end-of-month price prevailing just prior to the month of delisting.[10]

It is worth stressing the difference between the two insider portfolios in Eq. (1) and a portfolio based on equal-weighting firms and periods, where the latter is common in the literature on portfolio performance evaluation. The pervasive use of equal-weighted portfolios reflects a lack of granular data on investor holdings at less than the quarterly frequency. The holdings-based portfolio gives greater weight to firms and periods when insider holdings are high (measured by ω_{it}^{ow} and ω_{it}^{vw}). If insider holdings tend to be high due to price-sensitive inside information, then either of our two holdings-based portfolios have a better chance than the equal-weighted portfolio of identifying true insider performance. Holdings-based portfolio weights are also required to address the possibility suggested by Akbas, Jiang, and Koch (2020) that insiders with different personal investment horizons may view inside information differently and trade accordingly.[11]

[9]The value-weight ω_{it}^{vw} changes whenever firm i's stock price changes, even if insiders do not trade.

[10]As for the initial and final share-holdings of individuals (added and subtracted on the dates when when they became or ceased to be insiders according to our records), we follow the convention in the extant literature of not treating these as *bona fide* information-based purchases or sales.

[11]While stressing the advantage of using holdings-based portfolio weights throughout our analysis, we provide for comparison purposes a more standard performance analysis based on equal-weighted portfolios in Section 6 below.

3.2 Returns-based performance by gender

We estimate the performance of the two insider portfolios ($\omega_{it}^{ow}; \omega_{it}^{vw}$) classified by gender, and a third long-short (zero-investment) portfolio that is long in the male insider and short in the female insider portfolios, respectively. This estimation applies two returns-based methods of portfolio performance evaluation. Let $r_{pt}^e = r_{pt} - r_{ft} = \sum_{i=1}^{N_t} \omega_{it}^k (r_{it} - r_{ft})$ denote the return in month t to an insider portfolio with weights ω_{it}^k in excess of the risk-free rate r_{ft} (the monthly Norwegian Interbank Offered Rate or NIBOR). The two returns-based performance measures are measures of Jensen's α_{pt} defined as follows (where superscript 'hat' indicates OLS estimate):

$$\alpha_{pt} = \begin{cases} \alpha_{pt}^{4f} \equiv r_{pt}^e - [\widehat{\beta}_p^m (r_{mt} - r_{ft}) + \widehat{b}_p^{SMB} SMB_t + \widehat{b}_p^{HML} HML_t + \widehat{b}_p^{MOM} MOM_t] \\ \alpha_{pt}^{rb} \equiv r_{pt}^e - [\widehat{\beta}_{p,t-1}^{rb} (r_{mt} - r_{ft})] \end{cases} \tag{2}$$

The first performance metric is the constant term α_p^{4f} in a four-factor return regression (Fama and French, 1993; Carhart, 1997), where r_{mt} is the return on the equal-weighted market portfolio of OSE stocks, and the additional pricing factors SMB_t, HML_t and MOM_t are the returns to the FF-size factor (a portfolio of Small Minus Big stocks), the FF-value factor (a portfolio of High Minus Low book-to-market stocks) and the momentum factor (a long-short portfolio of stocks that is long in above-mean return and short in below-mean return over the past twelve months). All factors are generated within the OSE cross-section of stocks, much as in (Næs, Skjeltorp, and Ødegaard, 2008).

The second portfolio-based performance metric, α_{pt}^{rb}, is an estimate of the constant term in the rolling-beta estimation of the one-factor Capital Asset Pricing Model (CAPM), which allows for time variation in the portfolio's (lagged) market risk factor exposure $\beta_{p,t-1}^{rb}$. We report the average of these constant terms, $\alpha_p^{rb} = \frac{1}{T} \sum_{t=1}^{T} \alpha_{pt}^{rb}$. The estimate of the portfolio beta ($\widehat{\beta}_{p,t-1}^{rb}$) is calculated as a weighted average of beta estimates for the stocks in the portfolio: $\widehat{\beta}_{p,t-1}^{rb} = \sum_{i=1}^{N_t} \omega_{it} \widehat{\beta}_{i,t-1}$. For each firm i, the beta $\widehat{\beta}_{i,t-1}$ is estimated using three years of daily returns prior to the current month and the Scholes and Williams (1977) lead-lag beta adjustment for non-synchronous trading.

Table 3 summarizes the returns-based performance estimates for the portfolios of primary insiders. Panel A shows portfolio return descriptives, including average raw return, average excess return, and portfolio Sharpe Ratios calculated as mean$(r_p - r_f)$/sd$(r_p - r_f)$. For the long-short portfolios, the Sharpe ratio is mean(r_p)/sd(r_p). The Sharpe Ratio of the female insider portfolio with ownership weights

(Column 4) is 0.17, which is higher than the 0.09 for the male portfolio (Column 5). For the portfolio with insider value-weights, however, the Sharpe Ratio is similar across males and females: 0.11 and 0.10, respectively (columns 7 and 8).

Turning to the four-factor performance estimate in Panel B, notice first that the market exposures of the male and female portfolios tend to be similar. As expected for broad based portfolios, the market betas are all statistically significant and close to one. Female portfolios tend to have higher exposures than male portfolios to the two Fama-French size and B/M factors (*SMB* and *HML*). There is, however, no significant difference in male and female portfolio exposures to the momentum factor (*UMD*). As to the four-factor alphas in the first row of Panel B, α_p^{4f} for the Equal-weighted (buys) portfolio is .023 for females and 0.006 for males, respectively, which are significant at the 10% level.

Most important, none of the individual portfolio alpha-estimates are statistically significant at the 5% level or better. Moreover, the significance of the alphas of the long-short portfolios is even weaker. The lack of significance, and the consistently negative sign of the alphas of the long-short portfolios, clearly rejects the hypothesis that insider trades by males have better performance than those of females. This inference also holds when using the average rolling-beta estimation in Panel C of Table 3. Again, none of these recursive CAPM-alpha estimates, which allow for time variation in the estimated portfolio beta, are significant at the 5% level or better, nor are the alpha estimates of the long-short portfolios.

4 Holdings-based portfolio performance of primary insiders

Recall that the two returns-based performance metrics discussed above (α_p^{4f} and α_p^{rb}) measure average monthly abnormal portfolio returns. However, it is one thing to measure the average monthly performance of firms held by insiders, and quite another to measure the performance of the insiders' *actual* holdings. To illustrate, an insider that purchases additional shares just prior to the release of new and positive information about the company may not actually realize the positive stock price increase unless she also sells the stock by the end of the month. Weighting the monthly portfolio returns by the insiders' actual holdings therefore increases the power to detect whether insider truly buy low *and* sell high.

4.1 Holdings-based performance metrics

Insiders profit from private information by increasing the stock holding when future returns are likely to be higher than expected and reducing (not increasing) the holding when future returns are likely to be lower than anticipated by the market—resulting in a positive covariance between the change in the insiders' holdings and subsequent (abnormal) stock returns. Early applications of this covariance measure focused on individual investment portfolios (Cornell, 1979; Copeland and Mayers, 1982). Grinblatt and Titman (1993) and Ferson and Khang (2002) substantially expand the methodology and data to include broader samples of US mutual funds and actively managed investment portfolios, and Eckbo and Smith (1998) is the first to apply a conditional version of the covariance performance measure.

Let $r_{i,t+\tau}$ denote the realized return to firm i over the time horizon $t+\tau$. Our holdings-based performance measure (HM) is the covariance between the one-period change in the insider portfolio holdings from $t-1$ to t ($\Delta\omega_{it}$) and the τ-period future abnormal stock return, $r_{i,t+\tau} - E\left[r_{i,t+\tau}\right]$:

$$HM = \frac{1}{T-2}\sum_{t=2}^{T}\frac{1}{N_t}\left(\sum_{t=1}^{N_t} Cov\left(\Delta w_{it}, (r_{i,t+\tau} - E[r_{i,t+\tau}])\right)\right) \tag{3}$$

To compute the abnormal return, we first estimate $E[r_{i,t+\tau}]$ as the predicted return from a Fama-French-Carhart four factor model estimation. The four factor model is estimated in a similar manner to the earlier rolling beta CAPM estimation. That is, at date $t-1$ we estimate the four-factor model for each stock i using five years of data. This yields a (time varying) vector of coefficient estimates $\left\{\hat{\alpha}_{i,t-1}, \hat{\beta}^{m}_{i,t-1}, \hat{b}^{SMB}_{i,t-1}, \hat{b}^{HML}_{i,t-1}, \hat{b}^{MOM}_{i,t-1}\right\}$, which we use to generate an estimate of the the expected return $E[r_{i,t+\tau}]$.

We calculate HM for two alternative lagged benchmark portfolio weights ($\omega_{i,t-1}$):

$$\Delta\omega_{it} = \begin{cases} \omega^{ins}_{it} - \omega^{ins}_{i,t-1} & \text{lagged insider portfolio weights} \\ \omega^{ins}_{it} - \omega^{m}_{i,t-1} & \text{lagged OSE market portfolio weights} \end{cases} \tag{4}$$

where the second measure uses firm i's "CAPM-buy-and-hold" weight in the OSE market portfolio ($\omega^{m}_{i,t-1}$) as the benchmark for evaluation the change in insider portfolio weight (ω^{ins}_{i}) from $t-1$ to t. Moreover, we consider three alternative future return time horizons $(t+\tau)$:

$$\tau = \begin{cases} \text{1 month} & \text{short-lived insider information} \\ \text{3 months} & \text{intermediate-lived insider information} \\ \text{6 months} & \text{long-lived insider information} \end{cases} \tag{5}$$

We explore these three holding periods as HM will be positive only if the unobservable private inside information is made public and incorporated into the stock price during the period over which return is measured.

4.2 Holdings-based performance results

We test the null hypothesis that insiders do *not* trade on valuable inside information ($HM = 0$) against the alternative that they do ($HM > 0$). Table 4 reports the result of the estimation of HM, classified by gender. Note that, unlike for the returns-based tests in Table 3 above, it makes little sense to construct a long-short portfolio (male versus female). Rather, we test directly for equality of HM for male and female insiders and report the associated p-value in column (9) labeled p(diff).

There is no evidence in Table 4 that insider trades results in abnormal portfolio performance—that insiders know how to buy low and sell high. All of the values of HM are statistically insignificant at conventional levels. Consistent with our inference from the returns-based analysis above, the results in Table 4 fail to reject the hypothesis that the performances of male and female insiders are indistinguishable from zero and indistinguishable from each other. In conclusion, there is little evidence of a non-zero long-run abnormal performance of our insider portfolios, whether sorted on gender or not.

Finally, portfolio weights constructed based on decentralized trading decisions are, of course, mean-variance inefficient. For example, a central portfolio manager with access to valuable private information might choose to buy one security and sell another for hedging purposes (Grinblatt and Titman, 1989). While the degree of inefficiency of the decentralized insider portfolio weights is an open empirical question, Eckbo and Smith (1998) compare the performance of insider portfolios with that of the seven large mutual funds on the OSE over the period 1986–1992. They find that none of the portfolios they study are associated with abnormal performance, which suggests that the degree of inefficiency from decentralized trading decisions may be small.[12]

[12]Eckbo and Smith (1998) apply a conditional version of the covariance measure HM to a holdings-based portfolio of

We next turn to an event-study analysis of the short-term market reactions to insider trades. This helps answer the question of whether the market perceives insider trades to convey new information, which should result in abnormal firm performance in event-time within a few days of the trades.

5 Short-term performance around insider purchases

In this section, we estimate the short-term market reaction in event time around dates of insider purchases. The focus in this section is therefore on the average short-term performance of *firms* as they experience insider trades as opposed to the average performance of the *insiders* estimated above. We focus on insider purchases because the extant literature tends to conclude that stocks perform abnormally well following insider purchases, with negligible abnormal performance following insider sales (Cohen, Malloy, and Pomorski, 2012). Moreover, we screen out the routine trades explained in Section 2.2 above as these are unlikely to be based on inside information. Also, as the focus is on the immediate market reaction to the insider purchases, the analysis is performed using daily stock returns, which requires us to start the sample period after 1996 when insiders began reporting their trades within 24 hours.

We efficiently estimate the conditional abnormal return parameter γ_i in the following one-factor return-generating process for firm i:

$$r_{it}^e = a_i + b_i r_{mt}^e + \gamma_i D_{it} + \varepsilon_{it}, \quad t = t_{i1}, ...t_{i2}, \tag{6}$$

where r_{mt}^e is the return on the market portfolio in excess of the risk-free rate on day t. The dummy variable D_{it} takes a value of one inside an event window centered on the day of the insider purchase (day 0 in event time) and zero otherwise. We employ four alternative event windows: days (-1, 1), (-1, 5), (-1, 25), and (-1, 50). The start-date of the estimation, t_{i1}, is 1/1/1997 or, if later, the date firm i is first listed on the OSE. The end-date, t_{i2}, is the earlier of delisting and 12/31/2016. By construction, the event parameter γ_i measures the average daily abnormal return across all event windows experienced by firm i between t_{i1} and t_{i2}. This joint estimation of all firm-i events allows us to avoid the double-counting of overlapping event periods in calendar time that may otherwise occur when a series of events by the same firm are treated as independent in the estimation.[13]

insider trades on the OSE from the period 1985–1992.

[13]We have verified, however, that using the more standard residual-return approach (here with a fixed 250-day estimation

Ignoring for simplicity the firm-subscript i in Eq. (6), the cumulative abnormal return over firm i's k'th purchase event is $CAR_k = \tau_k\gamma_k$, where τ_k is the number of trading days in the k'th event window. Moreover, the t-statistic of CAR_k is $t_k = \tau_k\gamma_k/\sigma_{\tau\gamma} = \gamma_k/\sigma_\gamma$ where the standard deviation σ_γ is provided by the regression Eq. (6). If two event windows overlap in calendar time, we adjust (shorten) the second window so that they do not overlap, and adjust τ_k accordingly. Across a calendar period t_1 through t_2 with a total of K event windows, the resulting average cumulative abnormal return is therefore $CAR(K) = \sum_{k=1}^{K} \tau_k\gamma_k$.

Table 5 shows the results of the event study estimation. Interestingly, the average CAR is significantly positive for both male and female insiders over the two shortest windows, $(-1, 1)$ and $(-1, 5)$. Moreover, there is little if any difference in the CAR across gender. Thus, insider purchases convey positive firm-specific information to the market, whether the trades are by male or female insiders. It is possible that the mandatory insider-trade reports themselves convey some of this information directly to the market. While not tabulated, there is also a tendency for insider purchases to follow a multi-day negative abnormal return as if insider to some extent are able to "buy low" prior to a price rebound.

Evidence of a stock price rebound around insider trades is also reported in event studies on US data (Seyhun, 1986; Lakonishok and Lee, 2001; Jeng, Metrick, and Zeckhauser, 2003). Moreover, the evidence in Table 5 of gender-based short-run abnormal firm performance is somewhat comparable to that of Inci, Narayanan, and Seyhun (2017) on US data. However, the latter study also finds that male insiders tend to outperform their female counterparts, which is not the case in our data. We next exploit whether the Norwegian mandatory gender quota impacted the abnormal performance and the information content of primary female insiders.

6 Network effects of Norway's board gender-balancing

Recall from Figure 2 that, as a result of Norway's mandatory board gender balancing, the proportion of female directors in Norwegian OSE-listed companies rose from 15% in 2005 to 40% in 2008. The figure also shows that average board size has remained unchanged at five directors over the period 1998–2014.

period prior to the event and the exclusion of days with prior events in the estimation period) does not materially change our main conclusion below. See MacKinlay (1997) for a description of the standard residual-based approach to estimating event-induced abnormal returns, which treats multiple events by the same firm as independent. Thompson (1985) provides a general comparison of the conditional event-parameter estimation (such as in Eq. 6 above), while Kothari and Warner (2007) and Kolari and Pynnönnen (2010) discuss power issues in event studies.

This means that, rather than expanding board size to make room for new female directors (and keep existing male directors), firms typically chose to replace male directors with females.

In the most comprehensive economic study of this act to date, Eckbo, Nygaard, and Thorburn (2019) show that the director gender-quota law had a statistically insignificant impact on the market values of firms listed on the OSE. This evidence is reassuring from our perspective as it means that the shock to the population of insiders, while dramatically increasing the female director network, did not also systematically affect the stock returns that we use to measure insider performance. Thus, in this section, we examine whether the exogenous increase in the female director network has affected the absolute and relative performance of female primary insiders' trades.

Inci, Narayanan, and Seyhun (2017) suggest that female primary insiders in the US may be at an informational disadvantage relative to male directors due to the latter group's access to a much more extensive network of directors and executives. Building on this argument, we hypothesize that the greater female director network (both within firm and across firms) after 2007 has significantly enhanced primary female insiders' access to inside information. We test this hypothesis by contrasting estimates of abnormal trading performance before (1997–2007) and after (2008–2016) full compliance with the quota law.

Table 6 provide descriptive trading information for the two sub-periods 1997–2007 and 2008–2016 (using the same table format as Table 1 above). The fraction of the total number of insiders that are female increases from about 10% in the first subperiod to 22% in the second period. As expected, both the fraction of primary insider trades by females and the female insider transaction size increase substantially after the compliance deadline in 2007.

Tables 7 and 8 show the long-run primary insider performance using our returns-based and holdings-based performance measures, respectively, for the post-quota sample period 2008–2016. As reported above for the total sample period (tables 3 and 4), there is again no evidence of statistically significant abnormal performance for any of the portfolios or abnormal return measures, whether the insiders are male or female. Thus, we reject the hypothesis that the exogenous increase in female director network that resulted from the quota law has impacted insider trading performance.

For robustness, Table 9 reports the results of estimating returns-based abnormal performance of equal-weighted insider portfolios (weights are $1/N$ rather than as defined in Eq. 1). The portfolio formation, which follows closely that of Cohen, Malloy, and Pomorski (2012), is based on all (not just primary) insider trades and excludes "routine" trades as defined in Section 5 above. The abnormal return is estimated using the constant term of the four-factor Fama-French-Carhart regression model in Eq. (2). The estimation uses both the pre-quoa and post-quota time periods, 1997–2007 and 2008–2016, respectively. As shown, there is again no evidence of non-zero abnormal performance measured by the regression constant α_p^{4f} for either of the two subperiods.

Notwithstanding the absence of significant abnormal performance of the insider portfolios, an effect of the greater female network may have been to increase the information content of female insider purchases. The event-study estimates in Table 10, which is modeled on Table 5 above, suggest that this may indeed be the case. The table shows a statistically significant increase in the average $CAR(-1,1)$ and $CAR(-1,5)$ for female insiders from essentially zero in the pre-quota period (1997–2007) to a significant 0.15% and 0.14% in the post-quota period (2008–2016), respectively.

The increase in the CAR based on female purchases may reflect a combination of two effects, both driven by the board reform. First, the female network expansion may have given female insiders better access to valuable inside information, on which they trade occasionally. Second, the network expansion may have increased the stock market's confidence in the information conveyed by female insider trades. Note also that there is no evidence that the board reform has changed the CAR of male insiders: they realize a significant average CAR of similar magnitude in both subperiods. In sum, consistent with the estimates in Table 10, the board reform may have resulted in the average CAR of female insiders to become not only statistically significant but also of a magnitude that is now indistinguishable from the average CAR of male insiders.

In sum, while there is no evidence that the dramatic expansion of the female director network that happened leading up to year 2008 has had a measurable impact on the long-run holding-period performance of primary insiders trades (male or female), the short-term performance of female purchases has increased significantly, from zero to an average of 0.15% over the seven-day window following purchases. Again, given the absence of significant holdings-based abnormal portfolio performance, this short-term performance increase does not represent realized abnormal return to the insider. Rather, the increase in the short-term CAR is interesting because it suggests that the board reform may have increased the ability of female insiders to fine-tune their purchase orders and that outside investors appear to give more weight to the information in these reported trades.

7 Crisis trading and risk aversion

The above empirical analysis focuses on gender-based insider performance and on network information effects of Norway's mandatory board gender-balancing. In this section, we change the focus to insider trading activity observed during the financial crisis period 10/2008–12/2010. As we suggest below, differential shock-induced trading patterns between male and female insiders may be used to infer differences in risk aversion. As such, the trading-based evidence presented in this section contributes to the debate over gender-differences in risk aversion that go beyond the more conventional survey methods and experimental settings for comparing differences in risk aversion between male and female directors and executives (Adams and Funk, 2012).

7.1 Main hypothesis

As pointed out in the introduction, the stock market crash in September of 2008, which followed the bankruptcy of Lehman Brothers, likely had two effects on the incentive for insiders to purchase additional shares: (1) to take advantage of what insiders' considered temporary stock market underpricing, and (2) to restore an optimal share of equity in the individual insider's investment portfolio. Since the potential for underpricing was driven by an exogenous shock to the stock market, there is no particular reason to expect the mispricing motive *per se* to create a differential trading response among male and female insiders.

However, the stock-purchase demand created by (1) and (2) depends on the individual insider's risk aversion. Betting on a crash-induced market underpricing requires (temporarily) over-weighting the insider's stockholding in the firm where she is a primary insider. Moreover, restoring the optimal pre-crisis asset allocation between risky and risk-free investments requires increasing the weight of the market portfolio after the exogenous decline in the market. Risk aversion attenuates both these two trading motives. Thus, we hypothesize that, in the cross-section of individual insiders, less risk averse insiders buy more.

In the following, we examine this hypothesis by estimating trading propensities of primary insiders as a function of gender.

7.2 Director-level trading propensity

Figure 4 plots, at the firm level, the average fraction of female directors that trade in a given year, 1998–2014. The figure is based on the board data underlying Figure 2, which ends in 2014. Panel A shows female buy trades. The increase in 2006 and 2007 is likely driven by the incentives generated by the gender-quota law for newly appointed female directors to hold stocks in the firms they just joined as directors. This particular purchase effect expired by the end of 2007, when all OSE-listed firms were in full compliance with the 40% quota (Figure 2). Interestingly, Panels A and B of Figure 4 show that the purchase propensities of both female and male directors peaks in 2009—in the midst of the financial crisis. Moreover, this peak trading pattern is most dramatic for female directors. This is evidenced not only by the rate of increase in buy transactions in Panels A and B but also by the near-disappearance of sell orders in Panel C of Figure 2, which is unique to female directors.

We estimate the following probit model for the likelihood that an individual director j trades in quarter t, 1998–2014:

$$Y_{jt} = \alpha + \beta_1 Crisis_t + \beta_2 Control_{jt} + \epsilon_{jt}, \tag{7}$$

where the latent dependent variable $Y_{jt} = 1$ if the director trades and zero otherwise, and ϵ_{jt} is an error term. While $Y_{jt} = 1$ is given directly by our information on trades, we calculate the number of directors on each board that do *not* trade in the quarter ($Y_{jt} = 0$) as the difference between total (annual) board size and the number of trading directors. For example, if one director of a five-director board trades in quarter t, we add another four panel observations where $Y_{kt} = 0$ for that firm in quarter t, $k = 1, ..., 4$. Since we focus on possible gender differences in the trading likelihood, this way of constructing the data panel requires the assumption that the ratio of female to male board members is constant throughout the calendar year. This assumption is, however, easily defended as directors in Norway are elected for two-year terms.

Among the regressors, $Crisis_t$ is a dummy variable that takes a value of one during the 27-month crisis period 2008:10–2010:12. The vector $Control_{jt}$ contains characteristics of director j's firm that may affect the likelihood of a trade. These characteristics are $Market\ Cap$ (the natural log of the firm's market capitalization), $Volatility$ (the firm's quarterly stock return volatility) *Liquidity* (last quarter's average daily stock quoted bid/ask spread), and *Beta* (stock beta estimated over the past 36 months). These controls are included to capture the notion that it may be more difficult to trade based on price-sensitive information in larger, less opaque and more liquid stocks. Also, consistent with the information in Table 3 above, more risk averse directors may hold less and trade less the higher the systematic risk of the firm's stock. Table 3 shows that stock betas for portfolios held by female insiders are indeed lower than stock betas for portfolios held by male insiders. Finally, all regressions include industry fixed effects.

Table 11 shows the coefficient estimates for Eq. (7). In columns (1) and (2) we pool the male and female primary insiders and add the dummy variable $Female$ and the interaction term $Female * Crisis$ in order to test for gender-based trading differences. In columns (3)–(6) we run the regression separately for female and male trades. Since, over the sample period, male primary insiders on average hold a larger fraction of the firms shares than do female insiders, the coefficients on $Female$ in columns (1) and (2) are negative and significant. As expected, trading activity (both buys and sales) is higher for insiders in firms with higher liquidity (lower spreads) and volatility. Looking across to columns (3)–(6), this effect is more significant for male than for female insiders—likely due to the greater sample size. It suggests that insiders of both gender tend to concentrate their trades in less opaque firms.

Turning to the key variable of interest, the coefficients on $Crisis$ is large and statistically significant at the 1% level or better for purchases, and small and statistically insignificant for sales. Looking across the columns, both male and female directors increased their insider purchases significantly in response to the financial crisis. Also interesting, the coefficient estimate for $Crisis$ is similar in magnitude in columns (3) and (5), which is why the coefficient estimate on the interaction term $Female * Crisis$ is statistically insignificant in Column (1). In conclusion, Figure 4 and Table 11 show that female directors not only substantially increased their purchases during the financial crisis, but that they were also as likely as their male counterparts to purchase stocks. According to our main hypothesis above, this trading evidence suggests that female directors are no more risk averse than male directors.

7.3 Firm-level trading propensity of executives

We next use Eq. (7) to estimate the likelihood of an insider trade at the firm level, classified by whether the insider is a director or has an executive position in the firm. As we do not have information on the total number of executive positions in each firm, we perform the probit estimation at the firm level. Thus, the latent dependent variable is now Y_{it}, which takes a value of one if at least one insider in firm i trades in quarter t and zero otherwise. We are interested in executives' trades as these insiders are likely to be in a better position than board members to judge whether the firm's shares became (temporarily) undervalued during the financial crisis. Also, recall from Figure 3 above that executives drive much of the insider trades in general, including the increase in female insider trades during the crisis period.

The coefficient estimates for the trading propensity of executives are shown in Panel B of Table 12. The estimation employs the same control variables as in the director-level analysis of Table 11. Moreover, for comparison purposes, we re-estimate the propensity of director trades, this time at the firm level, shown in Panel A of the table. Beginning with the variables in the vector $Control_{it}$, for both female directors and executives, insider trading propensity (purchases and sales) is unrelated to stock return volatility and higher for larger and more liquid firms. Female executives are also more likely to purchase shares in larger and more liquid firms. For male insiders, firm-level executive (but not director) trading propensity also increase with firm size, while male insiders tend to trade more the more liquid the company.

Turning to the key variable of interest, *Crisis*, Table 12 shows that insider purchases (not sales) increases significantly for all types of insiders during the financial crisis period. That is, regardless of gender, and whether the insider is a director or an executive, the market decline caused by the financial crisis prompted all insiders to significantly step up their purchase activities. For our purposes, the important inference is that these trading data provide no basis for arguing that female insiders are more risk averse than their male counterparts.

7.4 Robustness

For purposes of robustness, we also estimate the effect of the financial crisis on the following two alternative measures of monthly aggregate insider trades, used previously by Lakonishok and Lee (2001) and Anginer, Donmez, Seyhun, and Zhang (2020):

$$Insider\ Direction_{it} = \frac{\sum_j Buy_{ijt} - \sum_j Sell_{ijt}}{\sum_j Buy_{ijt} + \sum_j Sell_{ijt}}, \tag{8}$$

where a Buy (Sell) is an indicator variable that takes a value of one if insider j in firm i has made a purchase (sale) in month t, and

$$Insider\ Shares_{it} = \frac{Shares\ Purchased_{it} - Shares\ Sold_{it}}{Shares\ Purchased_{it} + Shares\ Sold_{it}}, \tag{9}$$

where *Shares Purchased* $(Sold)_{it}$ is the total number of shares of company i purchased (sold) by insiders in month t. *Insider Direction* treats each insider trade equally, independent of the trade size, while *Insider Shares* gives more weight to larger trades in terms of the number of shares purchases or sold.

Figure 5 plots the fraction of companies at the OSE with positive *Insider Direction*, calculated separately for the trades of female and male insiders. The number of firms with a positive aggregate direction of inside trading clearly increases at the beginning of the crisis. This effect of the crisis is confirmed in Table 13, which reports coefficient estimates from panel regressions with either $Insider\ Direction_{it}$ or $Insider\ Shares_{it}$ as dependent variable. Again, the coefficient estimate for *Crisis* is positive and significant for both female and male insiders. Also as before, independent of gender, the coefficients indicate more insider trading in larger, more volatile, more liquid, and less risky firms.

8 Conclusion

This paper examines three decades of primary insider trades on the Oslo Stock Exchange—a stock market with an international reputation for being an insiders' market. In addition to testing for insider trading performance in general, our primary goal is to test for gender-based differences in insider holdings and trades by directors and executives. We construct holdings-weighted insider stock portfolios and subject these to an unprecedented range of econometric tests for long-run performance at a monthly frequency. Moreover, we report on the average short-term (daily) market reaction to insider trades, which is more commonly reported in the insider trading literature.

We classify all tests by insider gender, and subject each gender's trading activity to two important exogenous shocks. The first shock is the 2005 enactment of Norway's board gender-balancing legislation, which dramatically increased the female director network by forcing the percentage female directors from 15% to 40% in the course of two years. We examine whether this broad network increase improved female insiders' access to valuable inside information. The second shock is the financial crisis of 2008–2009. We use the relative female trading response to this exogenous decline in stock values to provide new evidence that challenges the conventional view that females tend to be more risk averse than males.

Beginning with the performance analysis, portfolios with weights constructed to reflect insiders' actual stock-holding show no evidence of abnormal insider performance at the monthly frequency, either over the total sample period (1998–2016) or in the period after gender-quota compliance (2008-2016). This conclusion is robust to varying performance metrics (both returns-based and portfolio holdings-based) and to varying the return horizon to capture long-lived inside information (up to six months following the trade date). Judging in particular from our novel holdings-based tests, which calculates the covariance between changes in insider stock holdings and subsequent abnormal stock returns, there is no evidence that insiders on the OSE succeed in 'buying low and selling high'.

Second, there is some evidence that, after passage of the gender-quota, female insiders are somewhat better able to time their stock purchases to days in which the firm realizes short-term (up to seven days) positive abnormal stock returns. We argue that a possible source of this increased timing is be incremental network information following the dramatic increase in the female director network caused by the gender quota. The informational network effect may have increased the market's valuation of the signal implied by female insider purchases. As for male insiders, there is evidence of short-term abnormal performance following purchases both before and after the forced gender balancing, with no change caused by the quota law.

The short-term firm-level abnormal return that we show exists following insider purchases is realized by the insider themselves only if they also sell their shares after the (cross-sectionally constant) short-term event window. Not surprising, our data on insider holdings show that this short holding period is literally never followed by insider sales. This fact explains why the short-term firm-level performance does not show up as abnormal insider return in our portfolios based directly on insiders' actual stock holding periods. It provides a general reminder to distinguish abnormal *firm* performance following insider purchases from the actual performance of the *insider*.

Finally, we provide robust evidence that female primary insiders—both directors and executives—significantly increased their purchases during the 27-month financial crisis period 2008:10–2010:12. Since it is unlikely that female insiders had better information than male insiders in terms of whether the crisis created a significant underpricing of their company's stock, the weight of this evidence points to a female director risk aversion that is no higher than male director risk aversion. This evidence is important as it is based not on surveys but directly on female insiders' investments, which are of substantial magnitude (the female director purchase transaction averages about sixty thousand U.S. dollars.) This evidence also supports the theory the selection of females as executives and directors may require these individuals to be 'more like men' in terms of risk aversion in order to break the glass ceiling.

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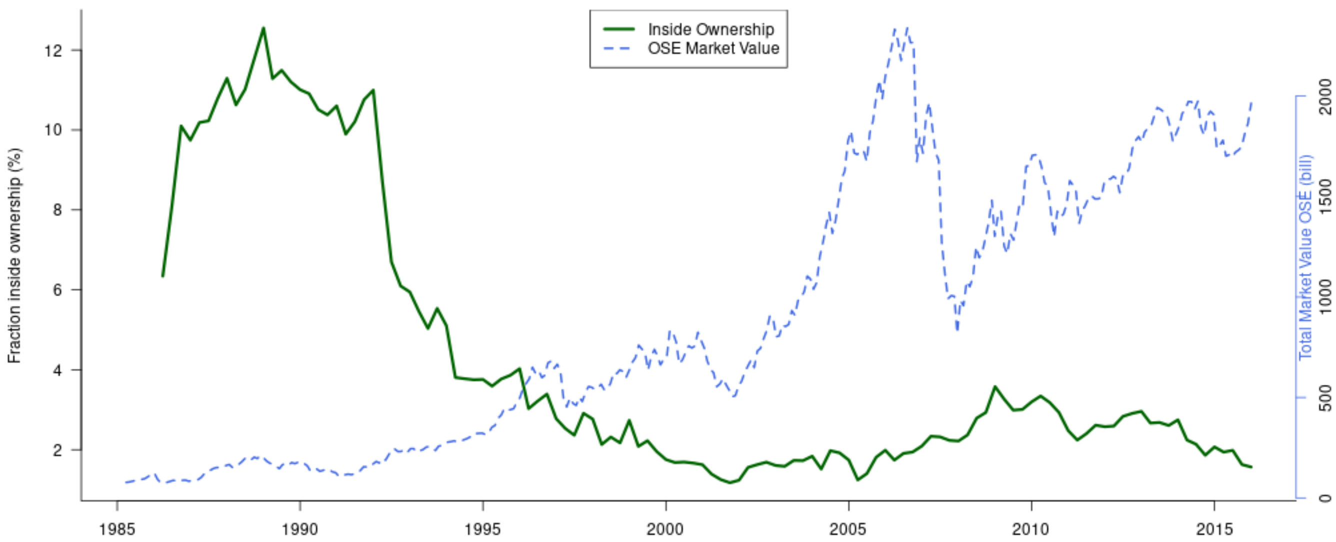
Figure 1: Fraction insider ownership on the Oslo Stock Exchange, 1998–2016

The inside ownership fraction is constructed for each company by summing the balances of all reporting insiders. The figures plot quarterly time series of cross-company averages of the inferred inside ownership fraction. The plot in panel A shows the average across all OSE companies. The plot in panel B shows the averages for companies grouped by market capitalization. The figure shows four groups, where group 4 contains the largest firms, and group 1 the smallest firms. Numbers are in percent. The plot in panel A also include the total market value of all OSE stocks, in billions, shown in the right-side vertical axis.

Panel A: All companies



Panel B: Companies sorted by size

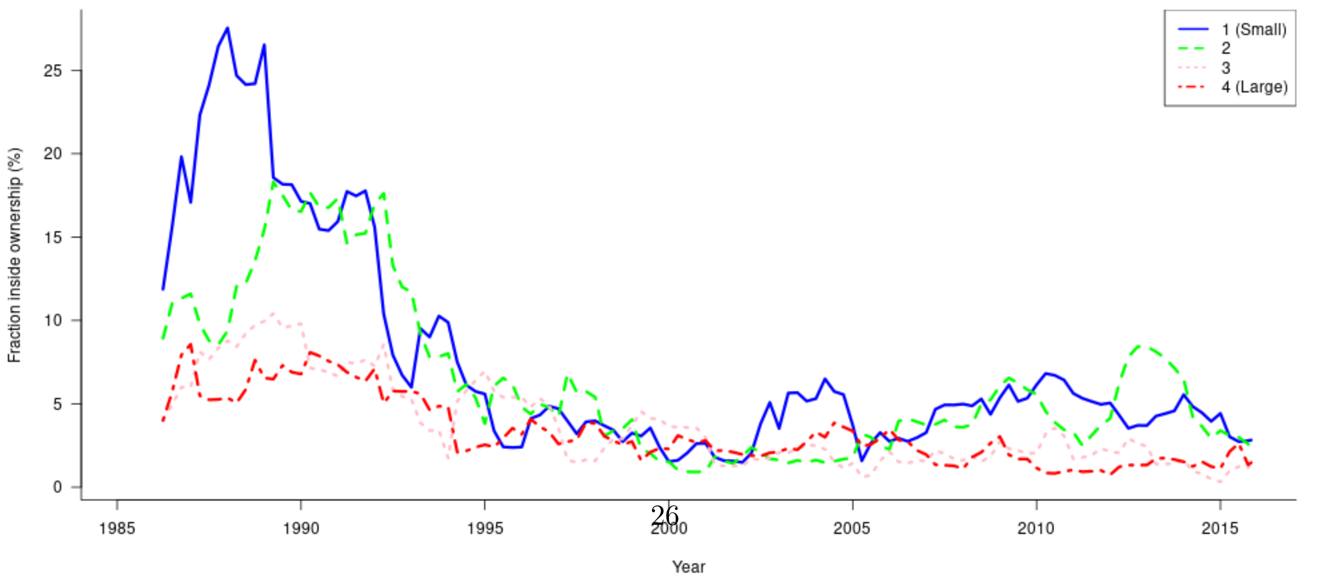


Figure 2: Board size and fraction female directors, OSE-listed firms 1998–2014.

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. Year 2008 (indicated with a vertical line) is the first year in which all Norwegian-domiciled OSE-listed firms are in full compliance with Norway's board gender-balancing law. Board data are from the national *Brønnøysund Registry Centre*, 1998-2014.

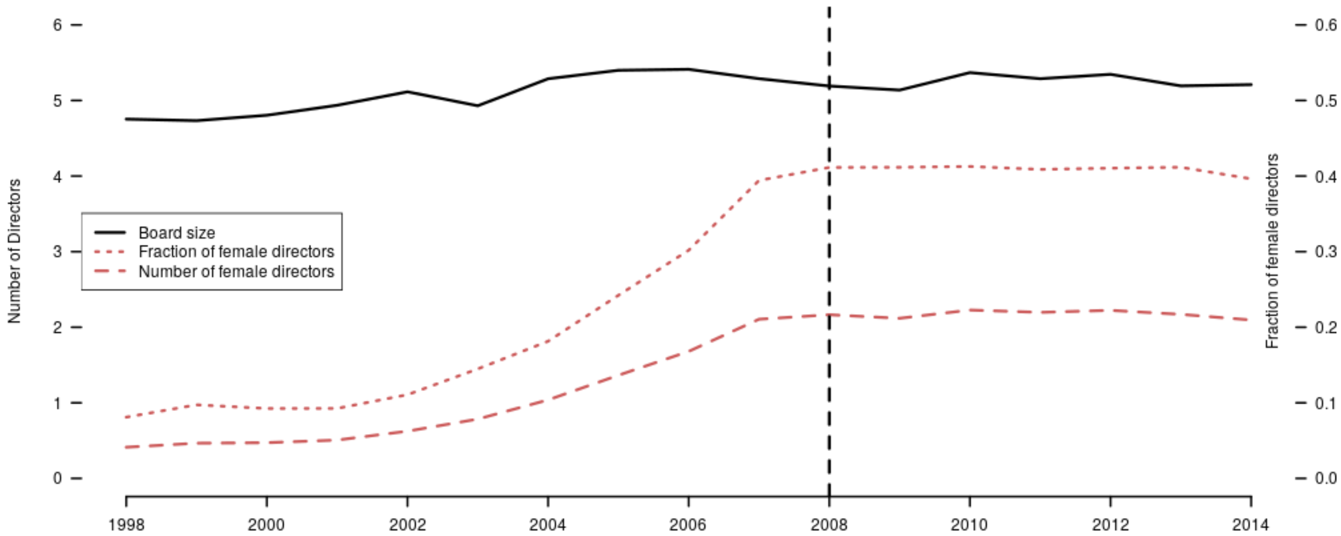


Figure 3: Percent of all primary insider trades by females, OSE 1986-2016

The figure plots the number of female primary insider trades in percent of all primary insider trades. Firms are listed on the Oslo Stock Exchange (OSE). Year-end 2007 marks the deadline for complying with Norway's board gender-balancing law, which was enacted in 2005 and require a minimum of approximately 40% of shareholder-elected directors to be from each gender.

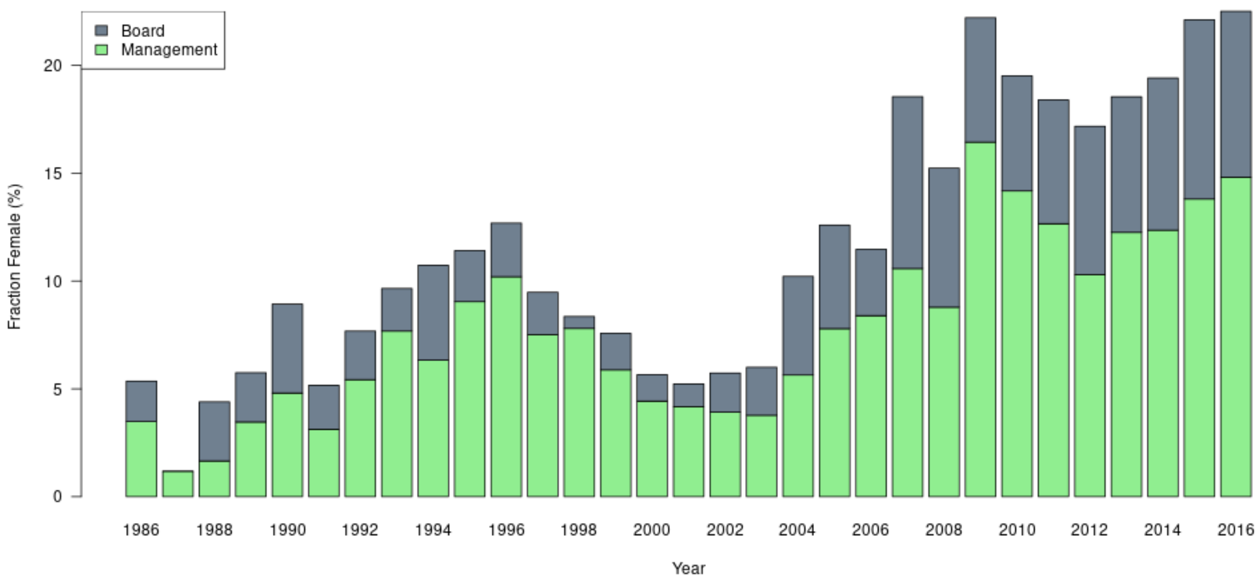
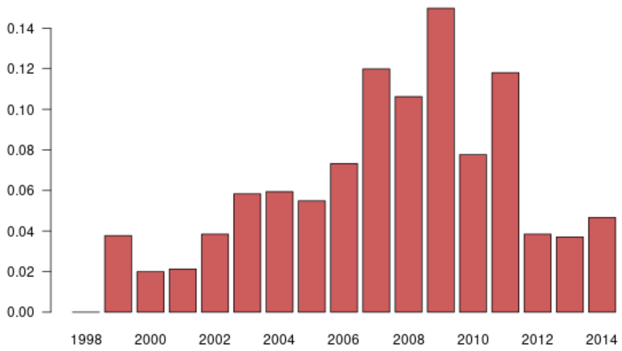


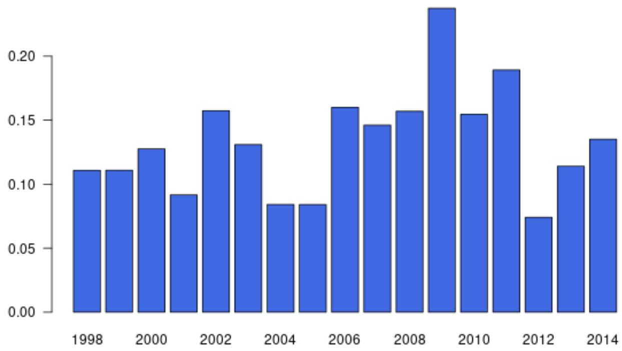
Figure 4: Fraction of male and female directors that trade, 1998–2014

The figure reports the annual average fraction of a board's directors, classified by gender, that report an insider purchase (panels A and B) or sale (panels C and D).

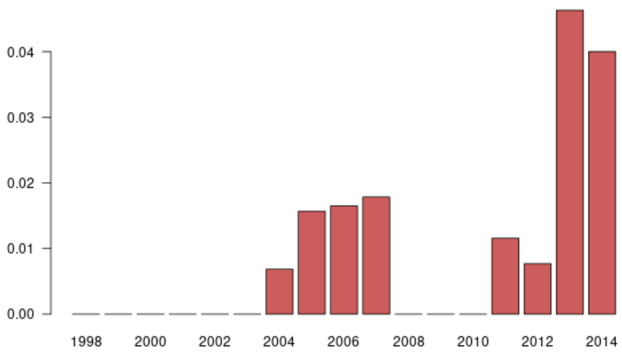
Panel A: Female buy trades



Panel B: Male buy trades



Panel C: Female sell trades



Panel D: Male sell trades

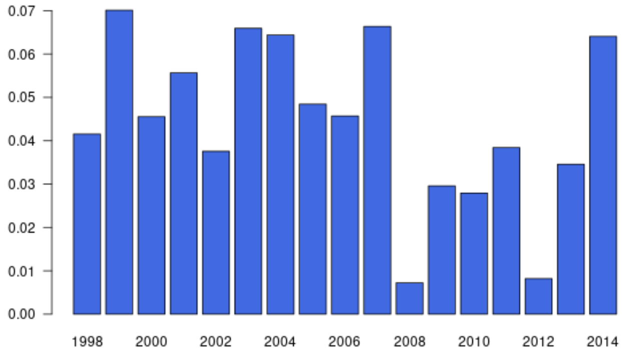


Figure 5: Fraction of positive *Insider direction*, 1997–2016

The figure plots the quarterly fractions of OSE-listed firms with positive aggregate *Insider Direction*, where

$$Insider\ Direction_{i,t} = \frac{\sum_j Buy_{ijt} - \sum_j Sell_{ijt}}{\sum_j Buy_{ijt} + \sum_j Sell_{ijt}}.$$

Buy (Sell) is an indicator variable that takes a value of one if insider j in firm i has made a purchase (sale) in quarter t.

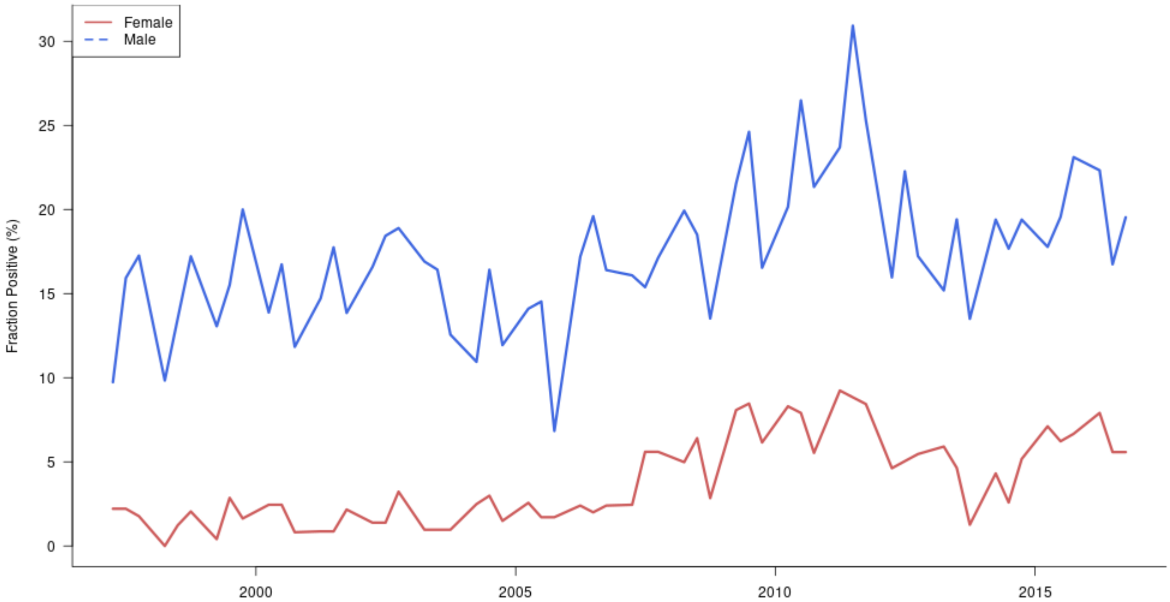


Table 1: Insider trading by gender: Sample descriptives

Primary insiders are directors and executives. Routine (repeated) trades are identified using the methodology of Cohen, Malloy, and Pomorski (2012), in which an insider trader in month t is classified as a "repeat performer" if the same insider traded in the same calendar month in each of the three years preceding the trade in month t. In Panel B, the number of distinct insiders is the number of individuals with insider transactions (excluding insiders who never transact). Panel C characterizes insider trading on an individual trader basis, using the insiders' trading history. The trading history begins with the first reported trade and ends with the last reported trade. We first compute the annual number of trades and trade values for each insider, and then form the sample period average for each insider (including years without trades). Panel B then reports the averages of these per insider averages. All value are in constant 2016 NOK using the consumer price index supplied by the Norwegian Bureau of Statistics (SSB).

A: Total sample of insider trades

	1986–2016		1986–1997		1997–2016	
	N	%	N	%	N	%
Total insider transaction records	47,429		23,213		24,223	
Records with *gender identified*	38,504	100%	17,098	100%	21,412	100%
of which by *primary insiders*	21,663	56%	5,660	33%	16,009	75%
of which are *non-routine*	19,108	88%	4,484	79%	14,630	91%

B: Transaction totals and averages

	All insiders				Primary Insiders			
	Total	Male	Female	Female(%)	Total	Male	Female	Female(%)
Number of firms	649	646	466	71.8	556	554	302	54.3
Number of distinct insiders	16473	13481	3003	18.2	5967	6100	1028	17.2
Total transaction value (mill.)								
Buys	161671	159945	1726	1.1	68628	67729	899	1.3
Sells	84902	82835	2067	2.4	69583	68341	1242	1.8
Number of transactions								
Buys	27082	23364	3718	13.7	16063	14387	1676	10.4
Sells	11422	10060	1362	11.9	5600	5195	405	7.2
Average transaction (1,000)								
Buys	5970	6846	464		4272	4708	536	
Sells	7433	8234	1518		12425	13155	3066	
Median transaction (1,000)								
Buys	74	88	27		119	130	62	
Sells	215	255	66		415	446	137	

C: Individual insiders' trading intensity

	All insiders			Primary insiders		
	All	Female	Male	All	Female	Male
Number of trades in year						
Buys	1.19	1.11	1.21	1.21	1.13	1.23
Sells	1.16	1.12	1.17	1.13	1.06	1.14
Annual transaction value (1,000)						
Buys	3839	461	4554	3026	631	3429
Sells	9795	1697	11204	17486	3297	19037

Table 2: Annual primary insider trades by gender, OSE 1986-2016

This table shows the annual distribution of the total of 21,663 primary insider trades from Table 1. Primary insiders are directors and executives. 100K means NOK 100.000.

	Primary insider purchases					**Primary insider sales**				
	Number of transactions				% Female	Number of transactions				% Female
Year	<100K	>100K	Female	Male	by value	<100K	>100K	Female	Male	by value
1986	19	44	3	60	0.01	17	27	2	42	0.13
1987	49	86	2	134	0.41	34	36	0	70	0.00
1988	55	91	7	139	0.09	37	34	2	71	0.03
1989	72	129	12	189	0.18	37	68	3	102	0.10
1990	123	224	24	323	0.16	127	106	20	214	1.12
1991	83	223	16	290	10.02	158	121	10	270	0.58
1992	112	254	17	349	1.01	100	94	0	194	0.00
1993	131	248	18	361	0.53	145	151	31	265	0.60
1994	169	321	35	457	1.79	129	117	30	218	4.69
1995	183	226	37	373	0.47	227	110	27	311	0.68
1996	248	345	59	535	0.79	192	146	22	316	0.18
1997	353	488	60	781	0.43	281	111	20	372	0.20
1998	187	230	21	398	0.05	87	28	7	108	0.10
1999	477	554	60	977	0.78	270	105	18	357	0.28
2000	277	270	20	529	0.15	218	34	13	239	14.24
2001	227	221	18	431	3.25	154	54	6	202	2.22
2002	261	286	24	523	0.12	69	43	3	109	0.01
2003	159	196	18	338	1.38	120	63	6	177	0.05
2004	149	168	25	294	0.26	123	38	15	146	0.59
2005	163	143	32	278	2.49	156	32	16	174	0.10
2006	306	156	41	424	0.32	223	26	15	235	0.69
2007	429	213	104	539	0.37	145	13	13	146	2.49
2008	345	275	84	538	7.31	61	15	3	73	0.04
2009	520	643	205	971	6.49	104	33	17	120	0.35
2010	487	531	162	866	14.50	98	31	14	115	3.00
2011	508	425	139	797	4.76	65	26	10	81	21.79
2012	314	191	66	440	1.45	80	24	17	87	14.17
2013	349	198	68	479	1.46	97	32	19	110	14.79
2014	402	247	91	559	2.77	96	35	20	111	25.61
2015	338	284	102	521	10.35	53	26	13	66	4.23
2016	295	302	106	494	10.37	69	38	13	94	2.39
All	7790	8212	1676	14387	1.44	3772	1817	405	5195	1.93

Table 3: Returns-based primary insider portfolio performance, 1986–2016.

The performance estimates reported in this table are based on monthly portfolio returns and rebalancing. The two sets of portfolio weights are defined in Eq. (4) in the text. The Insider-ownership-weight of firm i (columns 1-3) is the insiders' percentage ownership of firm i divided by the sum of the percentage insider holdings across all OSE firms. The Insider-value-weight (columns 4-6) of firm i is the value of insider holdings in i divided by the value of all insider holdings in all OSE firms. The Male–female portfolio is long in male and short in female insider weights, respectively. In Panel A, Sharpe Ratio is $\text{mean}(r_p - r_f)/\text{sd}(r_p - r_f)$ and, for the long-short portfolio, $\text{mean}(r_p)/\text{sd}(r_p)$. The two performance metrics, α_p^{4f} in Panel B and α_p^{rb} in Panel C, are defined in Eq. (2) in the text. The first is the constant term in a four-factor Fama-French-Carhart regression, while the second is the average constant term in a rolling-beta CAPM regression. Standard errors are in brackets, with p-values indicated as *=p<0.1, **=p<0.05, ***= p<0.01.

	Insider-ownership portfolio weights			Insider-value portfolio weights		
	Female (1)	Male (2)	Male–Female (3)	Female (4)	Male (5)	Male–Female (6)
A: Average raw returns and Sharpe Ratio						
$(1/N)\sum r_{pt}$	0.0198	0.0138	−0.0049	0.0136	0.0132	0.0012
$(1/N)\sum r_{pt}^e$	0.0165	0.0088		0.0086	0.0082	
Sharpe Ratio	0.1740	0.0929	−0.0537	0.0973	0.1066	0.0173
B: Four-factor alpha estimate						
α_p^{4f}	−0.0002	−0.001	−0.007	−0.003	0.00001	−0.0003
	(0.006)	(0.004)	(0.006)	(0.003)	(0.003)	(0.004)
β_p^m	0.897***	0.974***	0.215*	1.192***	1.109***	−0.089
	(0.108)	(0.082)	(0.113)	(0.065)	(0.049)	(0.070)
b_p^{SMB}	0.400***	0.072	−0.317**	0.091	−0.262***	−0.352***
	(0.126)	(0.103)	(0.132)	(0.081)	(0.062)	(0.088)
b_p^{HML}	0.273**	−0.150*	−0.387***	0.042	−0.187***	−0.256***
	(0.114)	(0.089)	(0.119)	(0.070)	(0.053)	(0.077)
b_p^{UMD}	0.074	0.149*	−0.141	0.001	0.066	0.068
	(0.094)	(0.076)	(0.098)	(0.060)	(0.045)	(0.065)
Observations	272	371	272	368	371	368
$\overline{R}^2$	0.211	0.275	0.073	0.488	0.611	0.057
C: Average rolling-beta CAPM estimate of alpha						
α_p^{rb}	0.0042	0.0011	−0.0079	0.0004	−0.0030	−0.0097*
	(0.005)	(0.004)	(0.005)	(0.005)	(0.004)	(0.005)
$\overline{\beta}_p^{rb}$	0.66	0.94	0.29	1.14	1.28	0.24

Table 4: Holdings-based primary insider performance, 1986–2016.

The holdings-based estimates in this table are based on covariances between monthly changes in insider holdings (weights) and subsequent returns, as follows:

$$HM = \frac{1}{T-2}\sum_{t=2}^{T}\frac{1}{N_t}\left(\sum_{t=1}^{N_t} Cov\left(\Delta w_{it},\ r_{i,t+T} - E[r_{i,t+T}]\right)\right)$$

where Δw_{it} is the change in the weight of stock i in the insider portfolio from month $t-1$ to t, and $r_{i,t+T} - E[r_{i,t+T}]$ is the abnormal returns over the subsequent T months ($T = 1, 3, 6$). Δw_{it} is either the monthly change in insider holdings, $w_{it}^{ins} - w_{i,t-1}^{ins}$, or the monthly change in insider holdings relative to the firm's weight in the OSE market portfolio (a CAPM "buy and hold" weight). $E[r_{i,t+T}]$ is the predicted return using the Fama-French-Carhart risk factors estimated using five years of monthly data prior to time t. The columns labelled p(diff) report the p-value for the test of difference between the male and female portfolio performance metrics. The p-values are calculated using standard errors that are robust to autocorrelation. Standard errors are in brackets, with p-values indicated as: *=p<0.1, **=p<0.05, ***= p<0.01.

	Insider-ownership portfolio weights			**Insider-value portfolio weights**		
Weight; Return	Female	Male	p(diff)	Female	Male	p(diff)
$Cov(\Delta w_{it};\ Ret_{i.t+T})$	(1)	(2)	(3)	(4)	(5)	(6)
A: Short-lived insider information: one-month future return horizon ($T = 1$)						
Δ_{it}: *lagged insider portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{ins}\ ;\ r_{i,t+1} - E[r_{i,t+1}])$	0.0012	−0.0008	0.26	0.0011	−0.0002	0.45
Δ_{it}: *market portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{m}\ ;\ r_{i,t+1} - E[r_{i,t+1}])$	0.0054	−0.0016	0.25	−0.0028	−0.0025	0.94
B; Intermediate-lived inside information: three-month future return horizon ($T = 3$)						
Δ_{it}: *lagged insider portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{ins}\ ;\ r_{i,t+3} - E[r_{i,t+3}])$	0.0025	−0.0002	0.19	0.0023	−0.0003	0.25
Δ_{it}: *market portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{m}\ ;\ r_{i,t+3} - E[r_{i,t+3}])$	0.0102	−0.0012	0.49	−0.0117	−0.0052	0.52
C: Long-lived insider information: six-month future return horizon ($T = 6$)						
Δ_{it}: *lagged insider portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{ins}\ ;\ r_{i,t+6} - E[r_{i,t+6}])$	0.0024	0.0002	0.27	0.0008	−0.0006	0.68
Δ_{it}: *market portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{m}\ ;\ r_{i,t+6} - E[r_{i,t+6}])$	0.0099	0.0004	0.88	−0.0232	−0.0077	0.38

Table 5: Market reaction to insider purchases, 1997–2016

The table reports the cumulative abnormal abnormal stock return $CAR \equiv \tau\gamma$, where γ is the average daily abnormal return over τ days in event time centered on the day of insider purchases (day 0) and estimated using the following one-factor return-generating process for firm i:

$$r^e_{it} = a_i + b_i r^e_{mt} + \gamma_i D_{it} + \varepsilon_{it},$$

where r^e_{mt} is the return on the market portfolio in excess of the risk-free rate on day t, and D_{it} is a dummy variable that takes a value of one inside the event window and zero otherwise. There are four alternative event windows around day 0: days $(-1,1)$, $(-1,5)$, $(-1,25)$, and $(-1,50)$. The estimation in Panel A (Panel B) uses trades of primary female (male) insiders only. We remove routine trades as in Cohen, Malloy, and Pomorski (2012). P-values are indicated by: *=p<0.1, **=p<0.05, ***=p<0.01.

Event windows:	$(-1,1)$	$(-1,5)$	$(-1,25)$	$(-1,50)$
A: Female Insiders				
CAR	0.012***	0.012***	0.007	0.003
	(0.001)	(0.0008)	(0.0004)	(0.0003)
Obs.	643,261	643,261	643,261	643,261
$\overline{R}^2$	0.030	0.030	0.030	0.030
B: Male Insiders				
CAR	0.014***	0.014***	0.003	−0.009
	(0.001)	(0.001)	(0.001)	(0.0004)
Obs.	1,013,513	1,013,513	1,013,513	1,013,513
$\overline{R}^2$	0.005	0.005	0.005	0.005

Table 6: Insider trades before and after quota compliance (1997–2007 v. 2008–2016)

The tables replicate the descriptives in Table 1 for two sub-periods: 1997–2007 and 2008–2016. In Panel A, the number of distinct insiders is the number of primary insiders with transactions (excluding insiders who never transact). Panel B characterizes insider trading on an individual trader basis, using the insiders' trading history. The trading history begins with the first reported trade and ends with the last reported trade. We first compute the annual number of trades and trade values for each insider, and then form the sample period average for each insider (including years without trades). Panel B then reports the averages of these per insider averages. All value as in constant 2016 kroner (NOK) using the consumer price index supplied by the Norwegian Bureau of Statistics (SSB). Data only for primary insiders.

A: Transaction totals and averages

	Primary Insiders							
	1997–2007				2008–2016			
	Total	Male	Female	Female(%)	Total	Male	Female	Female(%)
Number of distinct insiders	3394	3059	335	9.9	2913	2612	640	22.0
Total transaction value (million)								
Buys	45438	45229	208	0.5	9787	9261	526	5.4
Sells	54892	54235	657	1.2	5522	4995	528	9.6
Number of transactions								
Buys	5935	5512	423	7.1	6688	5665	1023	15.3
Sells	2397	2265	132	5.5	983	857	126	12.8
Average transaction (1,000)								
Buys	7656	8206	493		1463	1635	514	
Sells	22900	23945	4978		5618	5828	4187	
Median transaction (1,000)								
Buys	138	147	56		127	144	82	
Sells	742	793	168		505	611	195	

B: Individual insiders' trading frequency and intensity

	Primary insiders					
	1997–2007			2008–2016		
	All	Female	Male	All	Female	Male
Number of trades in year						
Buys	1.29	1.23	1.29	1.27	1.16	1.30
Sells	1.16	1.08	1.17	1.15	1.09	1.16
Annual transaction value (thousands)						
Buys	7529	698	8273	1725	613	1995
Sells	29795	6266	31652	5704	2107	6298

Table 7: Post-quota returns-based primary insider portfolio performance (2008–2016).

This table replicates the analysis in Table 3 for the period 2008–2016 after all OSE-listed companies were in full compliance with the quota law. The performance estimation uses monthly portfolio returns and rebalancing. The two sets of insider portfolio weights are defined in Eq. (4) in the text. The Insider-ownership-weight of firm i (columns 1-3) is the insiders' percentage ownership of firm i divided by the sum of the percentage insider holdings across all OSE firms. The Insider-value-weight (columns 4-6) of firm i is the value of insider holdings in i divided by the value of all insider holdings in all OSE firms. The Male–female portfolio is long in male and short in female insider weights, respectively. In Panel A, Sharpe Ratio is mean$(r_p - r_f)$/sd$(r_p - r_f)$ and, for the long-short portfolio, mean(r_p)/sd(r_p). The two performance metrics, α_p^{4f} in Panel B and α_p^{rb} in Panel C, are defined in Eq. (2) in the text. The first is the constant term in a four-factor Fama-French-Carhart regression, while the second is the average constant term in a rolling-beta CAPM regression. Standard errors are in brackets, with p-values indicated as *=p<0.1, **=p<0.05, ***= p<0.01

	Insider-ownership portfolio weights			**Insider-value portfolio weights**		
	Female (1)	Male (2)	Male−Female (3)	Female (4)	Male (5)	Male−Female (6)
A: Average raw returns and Sharpe Ratio						
$(1/N)\sum r_{pt}$	0.0080	0.0048	−0.0032	0.0102	0.0065	−0.0037
$(1/N)\sum r^e_{pt}$	0.0061	0.0029		0.0083	0.0046	
Sharpe Ratio	0.0603	0.0440	−0.0325	0.1376	0.0651	−0.0720
B: Four-factor alpha estimate						
α_p^{4f}	−0.004	−0.006	−0.004	−0.0004	−0.004	−0.006
	(0.010)	(0.004)	(0.010)	(0.004)	(0.004)	(0.005)
β_p^m	0.988***	1.170***	0.198	1.013***	1.096***	0.099
	(0.246)	(0.108)	(0.253)	(0.090)	(0.104)	(0.133)
b_p^{SMB}	0.153	0.057	−0.092	−0.277***	−0.474***	−0.193
	(0.240)	(0.106)	(0.247)	(0.088)	(0.101)	(0.130)
b_p^{HML}	0.045	−0.104	−0.147	−0.121	−0.076	0.046
	(0.247)	(0.109)	(0.254)	(0.090)	(0.104)	(0.134)
b_p^{UMD}	0.217	−0.050	−0.263	0.132*	0.129	0.001
	(0.189)	(0.083)	(0.195)	(0.069)	(0.080)	(0.102)
Observations	108	108	108	108	108	108
$\overline{R}^2$	0.113	0.595	0.008	0.669	0.677	0.007
C: Average rolling-beta CAPM estimate of alpha						
$\overline{\alpha}_p^{rb}$	−0.0004	−0.0040	−0.0055	−0.0034	−0.0060	−0.0046
	(0.009)	(0.004)	(0.009)	(0.009)	(0.005)	(0.009)
$\overline{\beta}_p^{rb}$	0.67	0.98	0.31	1.24	1.40	0.16

Table 8: Post-quota holdings-based primary insider performance (2008–2016)

The holdings-based estimates are based on covariances between monthly changes in insider holdings (weights) and subsequent returns, as follows:

$$HM = \frac{1}{T-2}\sum_{t=2}^{T}\frac{1}{N_t}\left(\sum_{t=1}^{N_t} Cov\left(\Delta w_{it},\ r_{i,t+T} - E[r_{i,t+T}]\right)\right)$$

where Δw_{it} is the change in the weight of stock i in the insider portfolio from month $t-1$ to t, and $r_{i,t+T} - E[r_{i,t+T}]$ is the abnormal returns over the subsequent T months ($T = 1, 3, 6$). Δw_{it} is either the monthly change in insider holdings, $w_{it}^{ins} - w_{i,t-1}^{ins}$, or the monthly change in insider holdings relative to the firm's weight in the OSE market portfolio w_{it}^{m} (a CAPM "buy and hold" weight). $E[r_{i,t+T}]$ is the predicted return using the Fama-French-Carhart risk factors estimated using five years of monthly data prior to time t. The columns labelled p(diff) report the p-value for the test of difference between the male and female portfolio performance metrics. The p-values are calculated using standard errors that are robust to autocorrelation. Standard errors are in brackets, with p-values indicated as: *=p<0.1, **=p<0.05, ***= p<0.01.

	Insider-ownership portfolio weights			**Insider-value portfolio weights**		
Weight; Return	Female	Male	p(diff)	Female	Male	p(diff)
$Cov(\Delta w_{it};\ Ret_{i.t+T})$	(1)	(2)	(3)	(4)	(5)	(6)
A: Short-lived insider information: one-month future return horizon ($T = 1$)						
Δ_{it}: *lagged insider portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{ins}\ ;\ r_{i,t+1} - E[r_{i,t+1}])$	0.0014	0.0004	0.70	0.0008	−0.0006	0.28
Δ_{it}: *market portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{m}\ ;\ r_{i,t+1} - E[r_{i,t+1}])$	0.0033	−0.0041	0.39	0.0008	0.0014	0.88
B; Intermediate-lived inside information: three-month future return horizon ($T = 3$)						
Δ_{it}: *lagged insider portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{ins}\ ;\ r_{i,t+1} - E[r_{i,t+1}])$	0.0018	0.0007	0.72	0.0009	−0.0024*	0.06
Δ_{it}: *market portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{m}\ ;\ r_{i,t+1} - E[r_{i,t+1}])$	0.0069	−0.0048	0.53	0.0004	0.0070	0.41
C: Long-lived insider information: six-month future return horizon ($T = 6$)						
Δ_{it}: *lagged insider portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{ins}\ ;\ r_{i,t+1} - E[r_{i,t+1}])$	0.0011	0.0012	1.00	0.0016	−0.0041*	0.09
Δ_{it}: *market portfolio weights*						
$Cov(w_{it}^{ins} - w_{i,t-1}^{m}\ ;\ r_{i,t+1} - E[r_{i,t+1}])$	0.0168	−0.0006	0.51	0.0058	0.0208	0.24

Table 9: Performance of equal-weighted insider portfolios before and after quota compliance

This table presents returns-based estimates of the abnormal performance of equal-weighted portfolios of insider purchase- and sale transactions, respectively, that follow closely the method in Cohen, Malloy, and Pomorski (2012). Thus, portfolios use all insider trades (not just primary) and are rebalanced monthly based on the previous month's trades. Moreover, routine trades are excluded using the algorithm of Cohen, Malloy, and Pomorski (2012) (discussed in Section 5 above). Also, the buy (sell) inside portfolio is constructed as a portfolio with at least one (non-routine) insider buy (sell) transaction in the previous month. Abnormal performance is estimated as the constant term α_p^{4f} in the four-factor Fama-French-Carhart regression model. Standard errors are in brackets, with p-values indicated as *=p<0.1, **=p<0.05, ***= p<0.01

	Buy Portfolio			**Sell Portfolio**		
	Female	Male	Male−Female	Female	Male	Male−Female
	(1)	(2)	(3)	(4)	(5)	(6)
A: Period before full gender-quota compliance, 1997–2007						
α_p^{4f}	0.005	0.006	−0.004	0.031***	0.014**	−0.021
	(0.008)	(0.004)	(0.009)	(0.011)	(0.006)	(0.013)
β_p^m	1.027***	1.335***	0.339**	0.840***	1.131***	0.160
	(0.158)	(0.069)	(0.170)	(0.227)	(0.116)	(0.263)
b_p^{SMB}	0.145	0.243**	0.112	0.102	0.203	0.119
	(0.210)	(0.094)	(0.226)	(0.284)	(0.158)	(0.329)
b_p^{HML}	−0.208	−0.037	0.180	−0.006	−0.170	−0.210
	(0.163)	(0.072)	(0.176)	(0.204)	(0.122)	(0.237)
b_p^{UMD}	0.239*	−0.221***	−0.486***	0.006	0.049	0.058
	(0.137)	(0.056)	(0.148)	(0.160)	(0.094)	(0.186)
Observations	122	132	122	97	132	97
$\overline{R}^2$	0.300	0.770	0.086	0.120	0.464	−0.017
B: Period after full gender-quota compliance, 2008–2016						
α_p^{4f}	0.025	0.004	−0.021	0.024	0.001	−0.018
	(0.019)	(0.006)	(0.016)	(0.027)	(0.008)	(0.026)
β_p^m	1.414***	1.618***	0.211	2.497***	1.047***	−1.553*
	(0.475)	(0.158)	(0.393)	(0.840)	(0.204)	(0.829)
b_p^{SMB}	0.378	0.196	−0.162	1.759***	−0.147	−1.788***
	(0.465)	(0.154)	(0.385)	(0.569)	(0.199)	(0.562)
b_p^{HML}	−0.300	−0.180	0.087	1.030	−0.102	−0.989
	(0.481)	(0.159)	(0.397)	(0.652)	(0.206)	(0.644)
b_p^{UMD}	−0.265	0.069	0.325	0.416	0.181	−0.244
	(0.365)	(0.121)	(0.302)	(0.523)	(0.158)	(0.516)
Observations	107	108	107	73	107	73
$\overline{R}^2$	0.078	0.542	−0.024	0.169	0.247	0.127

Table 10: Market reaction to insider purchases before and after quota compliance

The table reports the cumulative abnormal abnormal stock return $CAR \equiv \tau\gamma$, where γ is the average daily abnormal return over τ days in event time centered on the day of insider purchases (day 0) and estimated using the following one-factor return-generating process for firm i:

$$r^e_{it} = a_i + b_i r^e_{mt} + \gamma_i D_{it} + \varepsilon_{it},$$

where r^e_{mt} is the return on the market portfolio in excess of the risk-free rate on day t, and D_{it} is a dummy variable that takes a value of one inside the event window and zero otherwise. There are four alternative event windows around day 0: days $(-1, 1)$, $(-1, 5)$, $(-1, 25)$, and $(-1, 50)$. The estimation in panels A and C (panels B and D) uses trades of primary female (male) insiders only. In the estimation we remove routine trades as in Cohen, Malloy, and Pomorski (2012). P-values are indicated by: *=p<0.1, **=p<0.05, ***=p<0.01.

Event windows:	$(-1,1)$	$(-1,5)$	$(-1,25)$	$(-1,50)$
Prior to quota compliance: 1997–2007				
A: Female Insiders 1997–2007				
CAR	0.0026	0.0069	−0.0041	−0.0108
	(0.002)	(0.001)	(0.001)	(0.0004)
Obs.	209,427	209,427	209,427	209,427
$\overline{R}^2$	0.038	0.038	0.038	0.038
B: Male Insiders 1997–2007				
CAR	0.014***	0.014***	0.0064	−0.001
	(0.001)	(0.001)	(0.0003)	(0.0003)
n	507,385	507,385	507,385	507,385
$\overline{R}^2$	0.021	0.021	0.021	0.021
After quota compliance: 2008–2016				
C: Female Insiders 2008-2016				
CAR	0.0155***	0.0147***	0.0139	0.01126
	(0.001)	(0.001)	(0.001)	(0.0004)
Obs.	309,470	309,470	309,470	309,470
$\overline{R}^2$	0.027	0.027	0.027	0.027
D: Male Insiders 2008-2016				
CAR	0.014**	0.013	−0.0022	−0.0205
	(0.002)	(0.002)	(0.001)	(0.001)
Obs.	470,032	470,032	470,032	470,032
$\overline{R}^2$	0.003	0.003	0.003	0.003

Table 11: Effect of the financial crisis on director-level trading propensities

This table reports coefficient estimates in a probit model for the likelihood of an insider trade by an individual board member in a given quarter, 1998–2014. The variable *Female* is a indicator variable equal to one if the director is female. The indicator variable *Crisis* takes a value of one during the financial crisis period 2008:10–2010:12 and zero otherwise. The firm-level explanatory variables include *Market Cap* (the natural log of the firm's market capitalization), *Volatility* (the quarterly volatility of the firm's stock return), *Liquidity* (last quarter's average daily stock quoted bid/ask spread), and *Beta* (stock beta estimated over the past 36 months). The regressions include industry fixed effects for the 10 GICS industry codes. Statistical significance is indicated by p-values as follows: *=p<0.1, **=p<0.05, ***=p<0.01.

	All directors		**Female directors**		**Male directors**	
	Purchases	Sales	Purchases	Sales	Purchases	Sales
	(1)	(2)	(3)	(4)	(5)	(6)
Constant	−2.147***	−2.974***	−2.591***	−7.048***	−2.202***	−2.626***
	(0.225)	(0.341)	(0.617)	(1.674)	(0.243)	(0.355)
Crisis	0.191***	−0.076	0.230***	−4.151	0.194***	−0.086
	(0.048)	(0.085)	(0.086)	(279.138)	(0.048)	(0.085)
Female	−0.290***	−0.433***				
	(0.041)	(0.071)				
Female*Crisis	0.046	−3.008				
	(0.096)	(43.443)				
Market Cap	0.016*	0.035**	0.013	0.182***	0.020*	0.020
	(0.010)	(0.015)	(0.027)	(0.067)	(0.011)	(0.015)
Volatility	1.710***	1.443***	1.487*	9.517*	1.720***	1.403***
	(0.275)	(0.386)	(0.784)	(5.262)	(0.294)	(0.408)
Liquidity	−3.402***	−3.877***	−3.841**	−21.431*	−3.186***	−3.904***
	(0.573)	(0.950)	(1.628)	(11.500)	(0.613)	(0.969)
Beta	−0.049*	0.009	−0.038	−0.159	−0.048*	0.006
	(0.026)	(0.038)	(0.073)	(0.179)	(0.029)	(0.039)
Industry Fixed Effects	YES	YES	YES	YES	YES	YES
Observations	36,170	36,162	7,994	7,993	28,176	28,169

Table 12: Effect of the financial crisis on firm-level trading propensities

The table reports coefficient estimates in probit regressions of the likelihood of observing at least one insider trade in a given company. Estimated separately for gender and insider type using firm-quarter observations. In a given firm-quarter, the left-hand-side variable takes a value of one if there is an insider trade and zero otherwise. The table in Panel B report coefficient estimates of probit regression of the likelihood of an insider trade by an individual board member. The explanatory variables include the indicator variable *Crisis*, which takes a value of one during the financial crisis period 2008:10–2010:12. The firm-level explanatory variables include the log of the *market capitalization* of the firm, stock *volatility* (the quarterly volatility of the firm's stock return), stock *liquidity* (last quarter's average daily quoted stock bid/ask spread), stock *beta* (estimated over the past 36 months). The estimation period is 1998-2014. Statistical significance is indicated by p-values as follows: *=p<0.1, **=p<0.05, ***=p<0.01.

A: Directors

Insider trade among	**Female Directors**		**Male Directors**	
	Purchases	**Sales**	**Purchases**	**Sales**
	(1)	(2)	(3)	(4)
Constant	-3.176^{***}	-5.192^{***}	-1.413^{***}	-2.168^{***}
	(0.368)	(0.705)	(0.220)	(0.308)
Crisis	0.343^{***}	-0.262	0.136^{***}	-0.278^{***}
	(0.057)	(0.174)	(0.040)	(0.069)
Market Cap	0.061^{***}	0.113^{***}	0.007	0.019
	(0.016)	(0.030)	(0.010)	(0.014)
Volatility	0.559	0.763	1.006^{***}	0.823^{**}
	(0.481)	(0.570)	(0.329)	(0.361)
Liquidity	-3.418^{***}	-1.932	-2.084^{***}	-2.552^{***}
	(1.036)	(2.177)	(0.539)	(0.794)
Beta	-0.110^{***}	-0.036	-0.019	0.028
	(0.041)	(0.079)	(0.023)	(0.030)
Industry Fixed Effects	YES	YES	YES	YES
Observations	14,837	14,837	14,837	14,837

B: Executives

Insider trade among	**Female Executives**		**Male Executives**	
	Purchases	**Sales**	**Purchases**	**Sales**
	(1)	(2)	(3)	(4)
Constant	-3.850^{***}	-3.857^{***}	-2.128^{***}	-2.961^{***}
	(0.505)	(1.250)	(0.321)	(0.476)
Crisis	0.341^{***}	0.209	0.163^{***}	0.123^{*}
	(0.063)	(0.143)	(0.042)	(0.063)
Market Cap	0.091^{***}	0.092^{*}	0.055^{***}	0.061^{***}
	(0.022)	(0.054)	(0.014)	(0.021)
Volatility	0.745	-1.740	0.099	0.889^{*}
	(0.505)	(4.777)	(0.473)	(0.496)
Liquidity	-3.828^{***}	-14.930^{**}	-3.056^{***}	-2.288^{*}
	(1.480)	(6.397)	(0.857)	(1.284)
Beta	-0.062	-0.489^{***}	0.016	-0.078
	(0.056)	(0.161)	(0.035)	(0.054)
Industry Fixed Effects	YES	YES	YES	YES
Observations	5,791	5,791	5,791	5,791

Table 13: Effect of financial crisis on two alternative measures of insider trades

The table reports coefficient in cross-sectional regressions with the following two alternative measures of monthly aggregate insider trade as dependent variable:

$$Insider\ Direction_{i,t} = \frac{\sum_j Buy_{ijt} - \sum_j Sell_{ijt}}{\sum_j Buy_{ijt} + \sum_j Sell_{ijt}}$$

where Buy (Sell) is an indicator variable that takes a value of one if insider j in firm i has made a purchase (sale) in month t, and

$$Insider\ Shares_{it} = \frac{Shares\ Purchased_{it} - Shares\ Sold_{it}}{Shares\ Purchased_{it} + Shares\ Sold_{it}}$$

where $Shares\ Purchased\ (Sold)_{it}$ is the total number of shares of company i purchased (sold) by insiders in month t. The explanatory variables include the indicator variable *Crisis*, which takes a value of one during the financial crisis period 2008:10–2010:12. The firm-level explanatory variables include the log of the *market capitalization* of the firm, stock *volatility* (the quarterly volatility of the firm's stock return), stock *liquidity* (last quarter's average daily quoted stock bid/ask spread), and stock *beta* (estimated over the past 36 months). The estimation period is 1998-2014. Statistical significance is indicated by p-values as follows: *=p<0.1, **=p<0.05, ***=p<0.01.

Dependent variable:	*Insider Direction*		*Insider Shares*	
Sample gender:	Female	Male	Female	Male
	(1)	(2)	(3)	(4)
Constant	−0.059***	0.026	−0.058***	0.032
	(0.016)	(0.031)	(0.016)	(0.031)
Crisis	0.011***	0.012**	0.011***	0.010**
	(0.002)	(0.005)	(0.002)	(0.005)
Market Capitalization	0.004***	0.002	0.003***	0.002
	(0.001)	(0.001)	(0.001)	(0.001)
Volatility	0.067**	0.120**	0.067**	0.120**
	(0.029)	(0.055)	(0.029)	(0.056)
Liquidity	−0.088**	−0.273***	−0.092**	−0.264***
	(0.039)	(0.074)	(0.039)	(0.075)
Beta	−0.002	−0.003	−0.002	−0.003
	(0.002)	(0.003)	(0.002)	(0.003)
Industry Fixed Effects	YES	YES	YES	YES
Observations	24,143	24,143	24,143	24,143
$\overline{R}^2$	0.005	0.004	0.005	0.004