OPTION INITIATION AND UNDERLYING MARKET BEHAVIOR: EVIDENCE FROM NORWAY

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INTRODUCTION

Option introduction may have significant effects on the returns of the underlying security, as documented by a large number of studies on, primarily, U.S. stock markets. These effects are related to risk and return characteristics, to market microstructure, and to price adjustments to new information. Overall, option listing seems to have no impact on the systematic risk and a dampening effect on the total risk of the underlying stock. Furthermore, bid–ask spreads typically decline substantially, while the speed with which new information is incorporated into stock prices tends to increase. On the other hand, no firm conclusions can be drawn regarding the implications for stock excess returns and trading volume.

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This article sheds additional light on how option introduction affects underlying stock returns. It reports results from Norwegian markets. All four areas of research mentioned above are considered, and the findings are compared to the results of previous studies on foreign markets. The data indicate that option listing is associated with a temporary price increase, such that a positive excess return is available on the introduction day. Also, a sharp increase in trading volume and a substantial decline in bid–ask spreads are revealed. No significant change in either the systematic or the total risk of the stock can be found, however, an no improvements in the speed of price adjustments to new information can be discerned from the data.

Ex ante, one would expect the Norwegian evidence to be different from the results of studies on U.S. markets for several reasons. First, the Norwegian options and the underlying stock markets are small.[1] From a regulatory point of view, whether this fact yields features which are not seen in more significant markets is of special interest. Second, short selling of stocks is prohibited. The introduction of put options thus enables investors to trade on the basis of negative information much more effectively than was previously possible. Third, Norwegian options markets have existed for only a few years and cannot meet the maturity of the principal options markets. Fourth, small markets like Norway's may be more subject to manipulation, especially since Norwegian legislation does not explicitly prevent manipulation of options and stock prices. Finally, results from the Canadian markets are shown to be significantly different from broader U.S. and U.K. markets.

This article is organized as follows. In the next section, a brief review of previous studies on the effects of option listing is provided. In the third section, the data set of this study is explained. In the fourth section, the impact of option introduction on systematic and total risk is analyzed. Two regression models that adjust, respectively, for nonsynchronous trading and the time variation in volatility are utilized. In the fifth section, an event study is performed to explore any effects on excess stock returns. In the sixth section, the results from the two time-series regression models on the impact of option listing on trading volume and on bid–ask spreads are reported. In the seventh section, the adjustment of prices to new information is examined.

[1]Organized option trading was introduced at the Oslo Stock Exchange in 1990. Call and put options are currently listed on seven stocks. In 1993, option trading reached a total of about 1.28 million contracts, which was only 1/9 of the corresponding figure at the Stockholm Stock Exchange, and less than 1/16 of that for LIFFE.

PREVIOUS RESULTS

Following the Black–Scholes framework, most option pricing models replicate the payoffs from a derivative instrument using the underlying asset and riskless borrowing–lending opportunities. Thus, in this theoretical world an option is a redundant security, and option introduction can have no price effect on the underlying security. Outside this arbitrage-based approach in frictionless markets without transactions costs and taxes, price effects could exist for several reasons.

Trading in derivative assets may have a stabilizing effect on the underlying asset. Interaction between the spot and the derivative markets could result in more efficient risk allocation, which increases the demand for the underlying security and reduces spot price volatility. However, speculators may affect the information content of security prices negatively. If this effect dominates the risk pooling advantage, the underlying market becomes more volatile.

A large number of studies conclude that option listings seem to have no effect on beta risk, while total risk declines. Examples from U.S. markets are Damodaran and Lim (1991a), Haddad and Voorheis (1991), DeTemple and Jorion (1990), Skinner (1989), Bansal, Pruitt, and Wei (1989), Conrad (1989), and Ma and Rao (1986, 1988). Although early studies performed by e.g., Trennepohl and Dukes (1979) and Klemkosky and Maness (1980), did not support these findings, their results have been explained by some methodological deficiencies [see, e.g., Damodaran and Subrahmanyam (1992)]. The results of the U.K. study by Watt, Yadav, and Draper (1992) coincide with evidence from U.S. markets; but Chamberlain, Cheung, and Kwan (1993) could not find any significant change in either the systematic or the total risk of Canadian stocks.

Options may change the opportunity set available to investors by making an incomplete market more complete [see Ross (1976)]. Thus, the required rate of return could be reduced and cause an increase in the price of the underlying asset. Event studies on the price effect of option listing test for changes in excess stock return around the announcement or the introduction date. In an efficient market, one would expect that the announcement of option listing could have an immediate effect on stock prices. Rao and Ma (1987) found a negative excess return on announcement days. On the other hand, Conrad (1989) found no price impact around announcement dates, but found a permanent positive price change around option listing dates. This result is consistent with the findings in the early study by Branch and Finnerty (1981), which were later confirmed by DeTemple and Jorion (1990). Haddad and Voorheis (1991) reported a positive price effect only on introduction day, while Damodaran and Lim (1991a) identified a negative listing effect for put options where calls were previously listed. Furthermore, Watt, Yadav, and Draper (1992) were able to identify a temporary price increase from option listings in U.K. markets. Thus, no definite conclusions can be drawn with respect to the impact of option listings on excess returns.

Studies on microstructure effects of option introductions have focused on changes in stocks' trading volume and bid–ask spreads. Trading in options may result in a diversion of trading away from the underlying market to the options market, which results in a decrease in spot market liquidity. On the other hand, an increase in the size of the opportunity set makes risky investments more attractive, and, hence, option trading enhances the liquidity of the underlying market. Furthermore, an active options market may enable market makers in the underlying stock to hedge the risk of their portfolio inventory more effectively, which may lead to lower bid–ask spreads and greater liquidity in the underlying asset.

The empirical results are mixed. Some of the early studies on U.S. markets reported an increase in trading volume after the listing of options [see, e.g., Hayes and Tennenbaum (1979)]. Skinner (1989) and Damodaran and Lim (1991a) replicated this result using raw trading volume data. However, Damodaran and Lim found no significant change in market-adjusted volume. Bansal, Pruitt, and Wei (1989) categorized option listings by time period and documented increases in market-adjusted volume only prior to 1979. Chamberlain, Cheung, and Kwan (1993) could not reveal any significant changes in the Canadian markets, although a gradual increase in trading volume appeared to occur in the post-listing period on an unadjusted and on a market-adjusted basis.

Neal (1987) provided evidence that bid–ask spreads of the underlying stock decrease after option listing. Fedenia and Grammatikos (1989) and Damodaran and Lim (1991a) also identified a significant bid–ask spread reduction in the post-listing period. However, this was not true for the Canadian markets. Chamberlain, Cheung, and Kwan (1993) were unable to detect any significant reduction in bid–ask spreads.

The introduction of option trading may increase the quantity and improve the quality of information available about the underlying asset. Thus, the speed with which new information is incorporated into security prices may increase. A number of studies report such an effect. Jennings and Starks (1986) found that the prices of optioned stocks adjust more quickly to earnings reports than did their control sample of nonoptioned stocks. Damodaran and Lim (1991a) also documented that prices adjust more quickly to new information after option listing. Skinner (1990) reported a decline in excess returns following unexpected earnings news in the post-listing period. Watt, Yadav, and Draper (1992) provided evidence that the stocks adjust more rapidly to new information after option listing in the U.K.

Negative information can be more effectively revealed when put options are traded, especially when short selling is not allowed. Diamond and Verrecchia (1987) concluded that short selling restrictions imply that prices adjust more slowly to negative information. Thus, the introduction of put options is expected to increase the overall speed of price adjustments. Damodaran and Lim (1991b) found that prices adjust more quickly to negative earnings news after the listing of put options, while Watt, Yadav, and Draper (1992), including both calls and puts, reported a sharp decline in skewness for optioned stocks in the U.K.

DATA

Organized option trading was introduced at the Oslo Stock Exchange in 1990. Call and put series were introduced on five underlying securities for three- and six-month maturities, and on the OBX stock index for one-, two-, and four-month maturities. In September 1992, the market also opened for futures contract trading on the OBX stock index.

This study includes seven of the eight optioned stocks which have been listed by the Oslo Stock Exchange from its inception until November 1994. The excluded stock was introduced on the Exchange only a few weeks before its option listing date and therefore does not have the requisite 160 daily pre-listing returns. Table I reports the companies included in the test sample and the corresponding option listing dates. Hafslund Nycomed B was delisted after seven months of trading and replaced one year later by Hafslund Nycomed A.

The limited number of optioned stocks suggests that the results be compared with those of a control sample. For each option listing, a stock in a similar industry as the optioned stock and with a market value and trading volume of a similar magnitude to the optioned stock is selected for inclusion in the control sample. The names of the control stocks are also presented in Table I.

Daily prices for the stocks are from the Amadeus stock data base at the Norwegian School of Economics and Business Administration. All prices are adjusted for stock splits, dividends, etc. Volume data are from the Oslo Stock Exchange.

TABLE I
Sample of Optioned and Nonoptioned Stocks

Optioned Stocks *Company*	*Listing* *Date*	*Nonoptioned Stocks* *Company*
Bergesen d.y. B	May 22, 1990	Leif Høegh & Co.
Hafslund Nycomed B	May 22, 1990	Orkla A
Norsk Hydro	May 22, 1990	Nora A
Saga Petroleum A	May 22, 1990	Elektrisk Bureau
Hafslund Nycomed A	December 20, 1992	Dyno Industrier
Kværner A	April 29, 1993	Norske Skog A
Aker A	January 31, 1994	Rieber & Søn A

The sample of optioned stocks and their matched set constitute a major part of the Norwegian market. Analyzing risk and return characteristics, the Oslo Stock Exchange market value index and the individual returns would contain much of the same information. To circumvent this difficulty, a broader Scandinavian stock market index is utilized. It is a market value aggregate of the national stock market indices of Norway, Sweden, and Denmark. These data are from the Datastream data base and are provided by DnB Fonds. Over the period from January 1989 to October 1994, the average end of the month market value of the Oslo, Stockholm, and Copenhagen Stock Exchanges was 22,134, 92,070, and 31,392 million ecu, respectively. The Norwegian stocks account for only 15% of the total market value.

RISK

The systematic risk impact is tested by estimating market model regressions over 480 trading days. The data set is split into three subperiods covering days −320 to −161, −160 to −1, and +1 to +160, respectively, relative to the option listing day $t = 0$. When securities are thinly traded, OLS techniques yield biased beta estimates. To adjust for this problem, the approach of Fowler and Rorke (1983) [which was an extension of Scholes and Williams (1977)] is utilized here. Including two dummy variables, the following model is estimated:

$$r_{it} = \alpha_i + \beta_i^{-1} r_{m(t-1)} + \beta_i^0 r_{mt} + \beta_i^1 r_{m(t+1)} + c_i^1 D_i^1 r_{mt} + c_i^3 D_i^3 r_{mt} + e_{it} \tag{1}$$

where r_{it} is the day t return on security i; r_{mt} is the day t market return; D_i^1 and D_i^3 are dummy variables equal to unity in subperiods 1 and 3, respectively, and zero otherwise; and e_{it} is the error term. A c_i^1-coefficient significantly different from zero indicates an unstable beta in the pre-listing period, while a significant c_i^3-coefficient indicates that the option listing may have influenced the beta value. The null hypotheses that $\Sigma_i c_i^1 = 0$ and $\Sigma_i c_i^3 = 0$ are tested separately for both the test and the control sample. Also, a standard test for significant differences in beta between test and control companies is carried out.

The results are reported in Panel A of Table II. Although all average values of the coefficients are slightly negative, the null hypothesis cannot be rejected in any case (all p-values are beyond 11%). The beta values exhibit stability in the two subperiods before option listing. Option listing itself exhibits no statistically significant effect on average systematic risk. This is further confirmed by the p-value of 50.7%, when controlling for any significant differences in the c_i^3-value between the test and the control sample.

The revealed stability of betas between subperiod 1 and 2 also holds for the total risk of stock returns.[2] Consequently, all the following analyses are based on the 160 daily returns on either side of the option listing date. Changes in total risk from option listing are tested by estimating a standard market model in volatility over subperiods 2 and 3.

TABLE II
Stability of Beta and of Total Risk

Panel A: Estimated Stability Coefficients of Systematic Risk from Equation (1)

	Test Sample			*Control Sample*	
	c_i^1	c_i^3		c_i^1	c_i^3
Average c_i-value	−0.085	−0.124		−0.024	−0.049
p-value	0.274	0.114		0.789	0.541
Difference in the c_i^3-value			−0.075		
p-value			0.507		

Panel B: Estimated Stability Coefficients of Total Risk from Equation (2)

	Test Sample		*Control Sample*
Average g_i-value	−0.022		0.104
p-value	0.939		0.857
Difference in the g_i-value		−0.126	
p-value		0.843	

[2]The results are not reported here, but are available from the authors upon request.

Schwert (1989, 1990) and Schwert and Seguin (1990) account for any time variation in volatility by including the lagged stock and market return variance. Their approach is replicated here. In addition, the option listing is included as a dummy variable:

$$r_{it}^2 = a_i + b_i r_{i(t-1)}^2 + c_i r_{i(t-2)}^2 + d_i r_{mt}^2 + e_i r_{m(t-1)}^2 + f_i r_{m(t-2)}^2 + g_i D_i + e'_{it} \tag{2}$$

Since average returns are close to zero, r_{it}^2 and r_{mt}^2 are the measures used for the stock i and market return variance on day t.

The results are reported in Panel B of Table II. The average g_i-value is insignificantly negative for optioned stocks and insignificantly positive for nonoptioned stocks. The p-values are 93.9% and 85.7%, respectively. Therefore, no dampening effect of option listing on stock volatility in the Norwegian market can be established, which is further confirmed by an insignificant difference in the g_i-value between the two samples, where the p-value is 84.3%.

Option introduction is unlikely to have any impact on a company's business risk or capital structure. Stable beta values support this view. A reduction in total risk is consistent with the assumption that the introduction of options may change the opportunity set available to investors. DeTemple and Selden (1991) showed that when the market is incomplete, the introduction of an option may stabilize the market for the stock. Grossman (1988) pointed out that traded options, as opposed to synthetic options, reveal important information about future trading intentions of investors. The existence of such informationally relevant options may therefore lower the volatility of stock prices. On the other hand, introducing speculators may have a destabilizing effect on stock prices. Stein (1987, 1989) developed a model that showed that speculative agents affect the information content of prices adversely, thereby outweighing the price-reducing effects related to investors' pooling risk more efficiently.

The results of this study on total risk coincide with the Canadian experience. This may be due to shallower markets and a lower participation rate in the options markets than in the U.S. Furthermore, Skinner (1989) argued that the average reduction in return volatility increases by the number of stocks listed in a single market. This effect is the consequence of a gradual decrease in the size of the firms optioned, as the potential for volatility reduction in stock returns for smaller and less known companies is greater. Only a few stocks of the largest companies in Norway have been optioned, which may partly explain the insignificance of the decline in variances for Norwegian optioned stock returns.

EXCESS RETURN

An event study is performed to analyze the effect of option introduction on excess returns. The estimation period is 145 trading days before the start of the event period, i.e., $t = -160$ to -16, while the event window comprises 31 days, from day -15 to day $+15$. The null hypothesis is no unique impact on option contemporaneous stock prices. Thus, the expected value of average cumulative excess returns is zero.

To calculate excess returns, $AR_{i\tau}$, from the market model, Fowler and Rorke (1983) betas are used. Dodd and Warner (1983) propose a test statistic based on standardized excess returns, $SAR_{i\tau}$, which are calculated as follows:

$$SAR_{i\tau} = AR_{i\tau}/\sigma_{i\tau} \tag{3}$$

where $\sigma_{i\tau} = \{\sigma^2_{eit}[1 + 1/d + (r_{m\tau} - \bar{r}_m)^2/\Sigma_t(r_{mt} - \bar{r}_m)^2]\}^{1/2}$. σ^2_{eit} is the residual variance in company i's market model regression; d is the number of days used in the regression; $\bar{r}_m$ is the average market return in the estimation period; $r_{m\tau}$ is the market return on day τ (in the event window); and r_{mt} is the market return on day t (in the estimation period).

Table III reports both the average daily excess returns and the cumulative excess returns together with the associated p-values for the test and the control sample. Figure 1 depicts the cumulative excess returns for both groups. For expository purposes, the results for 10 days of each side of the option listing day are presented. On introduction day, optioned stocks are rewarded by a positive average excess return of 1.00%, which is significantly different from zero at the 5% level (the p-value is 0.8%). Since the control group displays only an insignificant positive average excess return of about 0.10%, the difference between the two samples is also significant (the p-value is 4.6%).

For the test group, cumulative average excess returns are mostly negative before and mostly positive after the introduction day. Comparing cumulative average excess returns for the test and the control sample, the excess returns are typically lower for the latter group from day -10 to day $+2$, and always higher thereafter. The high p-values associated with cumulative average excess returns suggest no apparent effect of option listing on excess returns besides the price increase on introduction day.

TABLE III
Average Excess and Cumulative Average Excess Returns

	Test Sample				Control Sample			
Day	Excess Return	p-value	Cumulative Excess Return	p-value	Excess Return	p-value	Cumulative Excess Return	p-value
−10	−0.323	0.393	−0.820	0.376	0.015	0.968	−1.418	0.126
−9	−0.267	0.480	−1.087	0.277	−0.365	0.334	−1.783	0.075
−8	−0.016	0.966	−1.103	0.302	0.458	0.226	−1.325	0.215
−7	0.200	0.597	−0.903	0.426	0.449	0.235	−0.876	0.440
−6	0.379	0.316	−0.524	0.661	0.569	0.132	−0.307	0.797
−5	0.441	0.243	−0.083	0.947	−0.779	0.040	−1.086	0.386
−4	0.030	0.937	−0.053	0.968	0.671	0.076	−0.415	0.751
−3	0.164	0.664	0.111	0.935	−0.156	0.680	−0.571	0.675
−2	−0.108	0.775	0.003	0.998	0.298	0.430	−0.273	0.847
−1	−0.392	0.300	−0.389	0.790	0.176	0.641	−0.097	0.947
0	1.001	0.008	0.612	0.686	0.098	0.795	0.001	0.999
1	−0.298	0.430	0.314	0.840	0.042	0.912	0.043	0.978
2	−0.513	0.175	−0.199	0.901	−0.467	0.212	−0.424	0.791
3	0.271	0.473	0.072	0.987	0.569	0.132	0.145	0.930
4	0.326	0.388	0.398	0.814	0.819	0.030	0.964	0.568
5	0.099	0.793	0.497	0.774	−0.241	0.524	0.723	0.676
6	−0.068	0.857	0.429	0.809	−0.065	0.863	0.658	0.711
7	−0.114	0.763	0.315	0.862	0.523	0.166	1.181	0.515
8	−0.204	0.589	0.111	0.952	−0.124	0.743	1.057	0.568
9	−0.055	0.884	0.056	0.976	−0.864	0.022	0.193	0.919
10	−0.221	0.559	−0.165	0.932	0.214	0.571	0.407	0.833

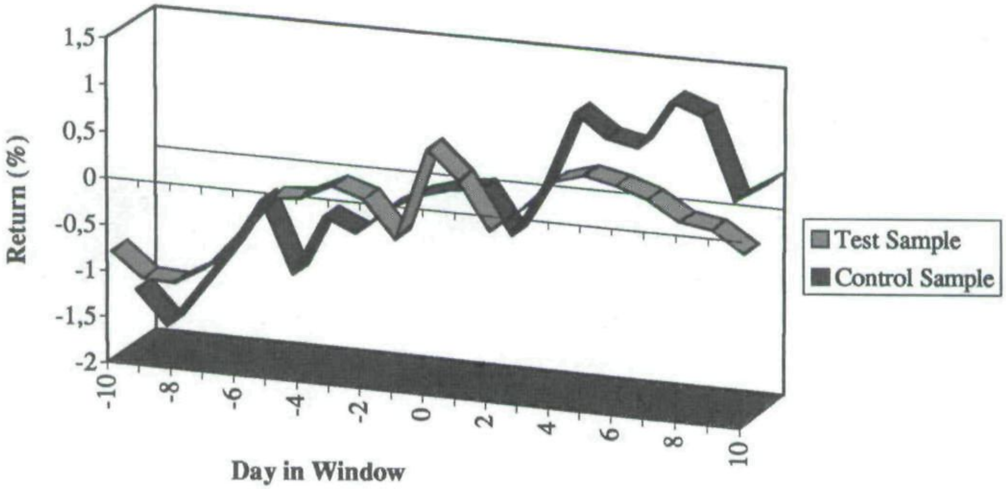

FIGURE 1
Cumulative average excess returns.

A temporary price change may be induced by liquidity suppliers who charge a premium for their services and by hedging-related demands of dealers. Conrad (1989) argued that the price effect may be sufficiently small, so that transactions costs make it unprofitable for any but those traders who anticipate acting as a dealer in the option and using the security to hedge.

Watt, Yadav, and Draper (1992) pointed out that this effect would be permanent if the options market makes an incomplete market more complete. An increased trading volume following option introduction could reduce the required rate of return on the stock, thereby also creating a permanent positive price effect. On the other hand, put options allow investors to get around short sales constraints more efficiently and permit negative information to be reflected in market prices. The negative listing effects of Damodaran and Lim (1991a) are consistent with this view. In Norway, call options were traded more actively than put options and this may explain the direction of the excess return.

A positive excess return only on listing day is consistent with the work of Haddad and Voorheis (1991). Their study contradicted earlier studies, i.e., Branch and Finnerty (1981) and Conrad (1989). Haddad and Voorheis concluded that the earlier mispricing may have been a transitory phenomenon resulting from the unfamiliarity of investors with the impact of option introduction on the underlying stocks, and that this disappeared when stock markets became more efficient as the options markets grew and matured. This is not an adequate explanation, however, in the present case.

MARKET MICROSTRUCTURE

Trading volume may be influenced by a number of factors related to option introduction. The net effect is not obvious, however. If trading volume is positively related to an increase in stock return volatility, see Schwert (1989), one would expect a decrease in volatility to be accompanied by a fall in trading volume. This might be the result of a diversion of trading away from the market for the stock to the market for its option [see Skinner (1989)]. On the other hand, options may increase the size of the opportunity set and thereby enhance the liquidity of the market. A greater institutional interest in the stock, greater media coverage, and closer scrutiny by analysts and investors may also increase trading volume and reduce bid–ask spreads [see Watt, Yadav, and Drapper (1992)].

The impact of option listings on trading volume in Norwegian stock markets is examined by estimating the following time-series regression model:

$$SV_{it} = a_i + b_i T + c_i D_i + \varepsilon_{it} \tag{4}$$

where SV_{it} is the market-adjusted turnover of stock i in month t; T is the time-index; and D_i is a dummy variable which is unity in the post-listing period and zero otherwise. SV_{it} is calculated by dividing the company's average monthly turnover by the total average monthly turnover at the Oslo Stock Exchange.

The results are reported in Panel A of Table IV. The trading volume for optioned stocks increased significantly. The p-value is 2.2%. Furthermore, an insignificant decrease (the p-value is 21.9%) occurs in the average c_i-coefficient for nonoptioned stocks, such that the difference between the test and the control sample is also significant. The p-value is 0.5%. In addition, if growth in trading volume is positively related to an increase in stock return variance, the reported results on total risk indicate that one would expect to find the higher trading volume in the control sample, rather than in the test sample.

Unlike the more recent studies, early U.S. studies reported a positive effect of option introduction on trading volume. The more mature the markets, the less effective an option listing seems to be in enhancing volume. Thus, observing an increased market-adjusted

TABLE IV

Stability of Trading Volume and of Bid–Ask Spread

Panel A: Estimated Stability Coefficients of Trading Volume from Equation (4)

	Test Sample		Control Sample
Average c_i-value	0.0184		−0.0035
p-value	0.022		0.219
Difference in the c_i-value		0.0219	
p-value		0.005	

Panel B: Estimated Stability Coefficients of Bid–Ask Spread from Equation (5)

	Test Sample		Control Sample
Average c_i-value	−0.00046		0.00225
p-value	0.184		0.108
Difference in the c_i-value		−0.00271	
p-value		0.030	

turnover in the post-listing period for Norwegian optioned stocks is not surprising. However, several factors influence volume and the result cannot be readily explained. Assuming a positive relationship between return volatility and volume, the negative impact on volume from the observed reduction in stock return variance in the post-listing period is outweighed by positive effects related to, e.g., an extended opportunity set for investors and improved hedging opportunities for market makers. In addition, if the largest companies have the lowest potential for return volatility reductions and their stocks are optioned first, the negative influence on trading volume from declined return variance would be fairly small in the starting period of organized option trading. Consequently, a positive net effect of option introduction is more likely to occur. This is at least consistent with the fact that the early studies on U.S. markets typically reported a positive relationship between option introduction and volume. The Canadian experience of no significant change in trading volume may be due to the fact that that study comprises a comparatively long period. The longer time span may have included some of the time when markets were already mature and this influenced the results of the study.

Arguments related to improved hedging opportunities for market makers also support the evidence of lower bid–ask spreads and greater liquidity. If the return volatility is reduced, the risk to the market maker of holding inventory will be lower and a smaller risk premium will be required. Furthermore, increased competition among market makers has a dampening effect on bid–ask spreads in the post-listing period. Finally, if option listing attracts informed traders away from the stocks, market makers will be less likely to be trading with investors better informed than themselves and thus be prepared to accept lower bid–ask spreads [see Chamberlain, Cheung, and Kwan (1993)].

To reveal any effects of option listing on the bid–ask spread, this time-series regression model is estimated:

$$SS_{it} = a_i + b_i T + c_i D_i + \varepsilon'_{it} \tag{5}$$

where SS_{it} is a standardized bid–ask spread for stock i on day t, T is the time index, and D_i is a dummy variable which is unity in the post-listing period and zero otherwise. SS_{it} is calculated by dividing the company's relative bid–ask spread by the market return variance over the specified time period, i.e., SS_{it} equals $\{(P_{bit} - P_{ait})/[(P_{bit} + P_{ait})/2]\}/\sigma^2_m$, where P_{bit} and P_{ait} are company i's bid and ask price on day t, respectively, and σ^2_m is the market return variance. Since the level of bid–ask spreads is naturally closely related to the volatility of Norwegian stock prices only, the Norwegian market index is utilized in the computation of standardized bid–ask spreads.

The results are given in Panel B of Table IV. The average c_i-value decreases insignificantly for the test companies (the p-value is 18.4%). However, the insignificant increase in the average c_i-coefficient for the control companies (the p-value is 10.8%) yields a significant difference between the c_i-estimates of the two samples (the p-value is 3.0%). Thus, option listings have a positive effect on the bid–ask spreads in the Norwegian markets.

ADJUSTMENT OF PRICES TO NEW INFORMATION

Low transactions costs and high leverage allow investors with private information to take positions in the options market more effectively than they are able to do in the underlying market. This fact is consistent with the documented increase in the speed of price adjustments to new information caused by option introduction. Furthermore, put options play an important role in revealing negative information to the market. This role is especially relevant in this study, since short selling is not allowed in Norway. In addition, when short selling constraints are made less severe, the unconditional distribution of price changes is expected to become less skewed to the left.

Before option listing, the average skewness of returns for the test sample is 1.03. In the post-listing period, it declines to 0.36. However, the control sample's skewness also declines, although a standard difference test cannot rule out that the improvement in skewness is identical for the two groups (the p-value is 45.3%).

The impact of option listing on the efficiency of price adjustment to new information is further examined by utilizing a simple model for price behavior, described by Amihud and Mendelson (1987):

$$P_{it} - P_{i(t-1)} = h_i(V_{it} - P_{i(t-1)}) + u_{it} \tag{6}$$

where P_{it} and V_{it} are logarithms of observed price and the intrinsic value of stock i, respectively, and h_i is a price adjustment coefficient for stock i. This model allows for both market structure-related and information noise, as well as imperfect price adjustments to value changes. Thus, information-related factors, e.g., trading-related mispricing, and market structure-related factors, e.g., bid–ask spreads, determine the error term u_{it}.

The observed variance of return can be decomposed into three components [see Damodaran and Lim (1991a)]:

$$\mathrm{Var}(r_{it}) = v_i^2 + 2\sigma_i^2 + \{[(2h_i - 2)/(2 - h_i)](v_i^2 + \sigma_i^2)\} \qquad (7)$$

where v_i^2 and σ_i^2 are the intrinsic variance of returns and the variance of the noise term, respectively; and the term in brackets is the contribution to variance from imperfect price adjustment. The first term is thus the variance portion attributed to the volatility of the underlying business. Underadjustment (overreaction) of prices to new information implies that $h_i < 1$ ($h_i > 1$), i.e., lower (higher) observed return variances.

To estimate price adjustment coefficients, the procedure described by Damodaran and Lim (1991a) is followed:

$$h_{ij} = [2\ \mathrm{Var}(r_{ijt}) - 2\sigma_i^2]/[\,jv_i^2 + \mathrm{Var}(r_{ijt})] \qquad (8)$$

where v_i^2 equals $(1/k)[\mathrm{Var}(r_{ikt}) + 2\ \mathrm{Cov}(r_{ikt,}r_{ik(t-1)})]$ and σ_i^2 equals $-\mathrm{Cov}(r_{ikt}, r_{ik(t-1)})$. For a given series of return data, only a value of k has to be estimated to obtain price adjustment factors over different intervals j. The results appear to be quite insensitive to the chosen value and, as in Damodaran and Lim (1991a), $k = 10$ is selected. A significant change in h_{ij} from the pre-listing to the post-listing period implies a change in the efficiency of price adjustment to new information.

According to the above-stated criterion, all stocks overreact to new information, i.e., $h_{ij} > 1$ in every case.[3] Surprisingly, this overreaction does not decline in the post-listing period for optioned stocks. In fact, the value of h_{i1} increases for all companies in the test sample. No distinct pattern is observed in the control sample.

The noise term of eq. (6) is influenced by mispricing caused by trading activities, which typically induces a positive first-order serial correlation in stock returns. This mispricing is corrected later when negative serial correlations in stock returns occur beyond lag 1. A possible effect of option listing can be related therefore to changes in the serial correlation structure of stock returns. For each company in the test and the control sample, the following regression model is estimated:

$$r_{it} = a_i + \Sigma_j b_j r_{i(t-j)} + \Sigma_j c_j D_j r_{i(t-j)} + u'_{it} \qquad (9)$$

where r_{it} is the day t return of stock i, and D_j is a dummy variable which is unity in the post-listing period and zero otherwise. One- through 15-day lags are calculated, i.e., $j = 1, \ldots, 15$.

[3]The results are not reported here, but are available from the authors upon request.

An F-test is performed to test the null hypothesis that all c_j-values equal zero. Of all stocks in both the test and the control sample, only one displays a significant change in the serial correlation structure (the p-value is 3.0%, while all the other p-values are beyond 25%). This is Hafslund Nycomed A, which was listed in the middle of the observation period. While no distinct pattern can be revealed for other stocks, a significant reduction from +8.56% to +1.95% in the first-order serial correlation is observed in this case. However, this indication of some growth in maturity in the Norwegian markets is not supported by the correlations results of the two more recent option introductions.

CONCLUSIONS

This article sheds more light on how option introduction affects underlying stock returns by examining the effects of option initiation in the Norwegian markets. Both public participants and regulatory authorities are concerned about unique effects that may arise in minor markets.

Recent studies on U.S. and U.K. markets document no change in systematic risk combined with a substantial decline in total risk. This study documents no change in systematic risk in the Norwegian markets, but also finds no significant change in total risk. This coincides with the Canadian experience. Since only stocks of the largest, well-known companies have been optioned in Norway, the potential for volatility reductions in stock returns is small.

Several studies on U.S. markets report a permanent positive price effect from option listing. The price increase may take place over several days or on introduction day only. In U.K. markets, a temporary price increase around option listing has been documented, which is confirmed in this study. However, the conclusion that a temporary price increase, as opposed to a permanent price increase, indicates more efficient stock markets as a result of growing and more mature options markets, cannot be supported. In Norway, call options are traded more actively than put options, which may indicate that investors do not utilize put options to get around short sales constraints. This fact may also account for the direction of the price change on introduction day.

In Norway, option introduction has a positive impact on stocks' trading volumes, which is the typical result for less mature markets. In these cases, the negative impact on trading volume from reduced stock volatility is small and outweighed by the positive effects from an extended investment opportunity set and improved hedging opportunities. A positive relationship exists between bid–ask spreads and trading volume.

Consequently, a substantial decline in bid–ask spreads in the Norwegian markets in the post-listing period is expected, and it is confirmed by the data.

Unlike U.S. and U.K. stock markets, no improved speed of price adjustment is observed after option listing. Furthermore, the ability to trade on the basis of negative information does not result in a statistically significant reduction of skewness of price changes. In addition, only one stock displays a substantial decline in the serial correlation structure of returns in the post-listing period.

The experience from Norwegian option introductions is generally positive. The results coincide largely with those from foreign markets and demonstrate that the regulatory authorities do not have to be too concerned about unique effects arising from option introductions in these minor markets. However, unlike most previous studies, the data reveal no improvements in the speed of stock price adjustments to new information. This result indicates that imposing more stringent regulations to prevent manipulation of prices could further increase the positive impact of option listing on the underlying stock market.

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