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Mikko Mäkinen

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ABSTRACT: The essays in this thesis study stock option schemes and CEO compensation in the publicly listed Finnish firms. The first essay studies the determinants of option scheme adoption. It argues that firms with a higher level of market value per employee are able to provide economic incentives for their personnel at a lower cost, thus encouraging the use of options. In addition, share returns from the past year affect the adoption of selective option schemes, but not broad-based plans. The second essay presents empirical evidence on the productivity impacts of stock option schemes. It shows that when endogeneity and dynamics of options are taken into account, there is no statistical evidence on the link between option schemes and firm productivity. The finding is consistent with the hypotheses that predict negligible effects of option plans on firm performance. The third essay studies whether, instead of productivity, option schemes affect firm technical inefficiency. The findings indicate that in the manufacturing sector broad-based scheme firms have higher mean inefficiency than selective or non-option firms. There is also no statistical support for the hypothesis that option schemes may reduce firm technical inefficiency. The fourth essay studies statistical relationships between CEO compensation, firm performance and firm size. It shows that CEO average compensation has increased substantially in 1996-2002. However, the change in CEO compensation, and especially in the total compensation, can be associated with changes in stock market-based measures of firm performance, such as shareholder value and share return. Also, changes in both accounting (ROA%) and stock market-based firm performance measures in the previous year can be associated with the change in CEO compensation in the following year. CEO pay level is significantly related to firm size. Typically, pay-for-firm size elasticity is in the range of 0.2-0.3.

Key words: stock options, productivity, technical inefficiency, ceo compensation

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TIIVISTELMÄ: Tutkimuksessa tarkastellaan optio-ohjelmia ja toimitusjohtajan palkkausta julkisesti noteeratuissa suomalaisyhtiöissä. Ensimmäisessä esseessä tutkitaan optio-ohjelmien käyttöönottoon vaikuttavia tekijöitä. Esseessä esitetään, että yritykset, joiden markkina-arvo per työntekijä on korkea, voivat tarjota taloudellisia kannustimia henkilöstölleen edullisesti, mikä kannustaa yrityksiä käyttämään optioita. Lisäksi yhtiön edellisen vuoden osaketuotto vaikuttaa kohdennettujen mutta ei laajapohjaisten optio-ohjelmien käyttöönottoon. Toisessa esseessä esitetään empiiristä evidenssiä optio-ohjelmien tuottavuusvaikutuksista. Esseessä osoitetaan, että jos optio-ohjelmien endogeenisuus ja dynaamiset vaikutukset huomioidaan, optioiden ja yrityksen tuottavuuden välillä ei löydy tilastollisesti merkitsevää yhteyttä. Havainto tukee hypoteesia, jonka mukaan optio-ohjelmien vaikutus yrityksen taloudellisen menestymiseen on vähäinen. Kolmannessa esseessä tarkastellaan vaikuttavatko optio-ohjelmat, tuottavuuden sijaan, yrityksen tekniseen tehottomuuteen. Tulokset viittaavat siihen, että teollisuusyrityksissä, joissa on käytössä laajapohjainen optio-ohjelma, tekninen tehottomuus on keskimäärin korkeampi kuin teollisuusyrityksissä, joissa on kohdennettu optiojärjestelmä tai ei lain-

kaan järjestelmää. Tulokset eivät tue hypoteesia, jonka mukaan optio-ohjelmat vähentävät yrityksen teknistä tehottomuutta. Neljännessä esseessä tarkastellaan tilastollista yhteyttä toimitusjohtajan palkan sekä hänen johtamansa yrityksen taloudellisen menestymisen ja koon välillä. Esseessä osoitetaan, että toimitusjohtajan keskimääräinen palkka on noussut tuntuvasti ajanjaksolla 1996-2002. Toimitusjohtajan peruspalkan muutos, ja erityisesti kokonaispalkan muutos, voidaan yhdistää hänen johtamansa yhtiön taloudelliseen menestymiseen, kun yhtiön menestymistä mitataan osakemarkkinapohjaisilla mittareilla, kuten omistaja-arvolla ja osaketuotolla. Lisäksi toimitusjohtajan palkan muutos voidaan yhdistää hänen johtamansa yhtiön edellisen vuoden taloudelliseen menestymiseen, kun yhtiön menestymistä mitataan kokonaispääoman tuotolla (ROA%) tai osakemarkkinapohjaisilla mittareilla. Toimitusjohtajan palkkataso riippuu merkitsevästi yrityksen koosta. Tyypillisesti estimoitu palkkajousto yrityksen koon suhteen on noin 0.2-0.3.

Asiasanat: työsuhdeoptiot, tuottavuus, tekninen tehottomuus, toimitusjohtajan palkkaus

PREFACE

This study consists of an introductory essay and the following four empirical essays on stock option schemes and CEO compensation in Finland:

1. The Determinants of Stock Option Compensation: Evidence from Finland. *Industrial Relations, vol. 45, No. 3 (July 2006).* (Joint essay with Derek C. Jones and Panu Kalmi).

2. The Productivity Effects of Stock Option Schemes: Evidence from Finnish Panel Data. *Unpublished manuscript.* (Joint essay with Derek C. Jones and Panu Kalmi).

3. Do Stock Option Schemes Affect Firm Technical Inefficiency? Evidence from Finland. *Unpublished manuscript.*

4. CEO Compensation, Firm Size and Firm Performance: Evidence from Finnish Panel Data. *Unpublished manuscript.*

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Two of the essays are joint studies with Professor Derek C. Jones and Doctor Panu Kalmi. Working with Derek and Panu has not only been a great privilege but also educating and highly pleasant. I am deeply grateful to both of you.

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Helsinki, January 2007

Mikko Mäkinen

ESSAYS ON STOCK OPTION SCHEMES AND CEO COMPENSATION: AN INTRODUCTORY ESSAY

In recent years, the popularity of stock option schemes and the increases in CEO compensation have generated widespread public debate in several countries. The four essays of this study aim to provide new evidence on these issues in Finland. The essays can be categorized by their key research questions as follows: Essay 1: How do firm characteristics and stock market conditions affect option programs' adoption? Essay 2: What is the impact of option schemes on firm productivity? Essay 3: Do option schemes affect firm technical inefficiency? Essay 4: What is the relationship between CEO compensation, firm size and firm performance?

I next describe the related literature and review the essays in the collection in more detail.

Related Literature and Essays

The first essay analyzes several hypotheses that may explain the adoption of selective and broad-based stock option programs. Core and Guay (2001), Ittner, Lambert and Larcker (2003), and Kroumova and Sesil (forthcoming) highlight human capital intensity in a firm's production process as an important determinant of the adoption of option plans. When work is human capital intensive, it becomes difficult to monitor and, in lieu of direct supervision, employees need self-motivating incentive schemes. However, accounting-based group incentive schemes may be problematical, since it is often difficult to value intangibles correctly in R&D intensive organizations, and therefore, market-based compensation seems to be preferred in these situations.

Bergman and Jenter (2004) develop an elaborate argument to try to account for the puzzling observation that stock options are used at all to compensate risk-averse employees. They argue that it is necessary that risk-averse employees have more optimistic expectations concerning stock price developments than do outside risk-neutral investors. They attribute this optimism to "excessive extrapolation", where employees form expectations based on past share returns and believe that high returns will continue in the future.

A standard prediction from the principal-agent theory suggests that risk-averse employees would dislike schemes where a part of their pay is tied to a volatile incentive component, and where a significant part of the volatility is beyond the control of employees. However, Prendergast (2002) has perceived that empirical research often finds a positive relation between risk and incentives, since in more uncertain settings the principal is often better off delegating responsibility to the agent(s). Oyer (2004) argues that when uncertainty is high, fixed wage contracts require frequent revision, but the transaction costs of rewriting the contracts become prohibitively costly. To retain the best employees, it is better to tie compensation to a measure that correlates with the business cycle. Again this gives rise to a positive correlation between risk and option compensation, which Oyer and Schaefer (2005) observe in their empirical analysis.

The famous 1/n problem (Alchian and Demsetz, 1972) suggests that group-based incentive schemes, such as equity pay, become ineffective when the size of the group grows. According to Alchian and Demsetz, larger groups necessitate hierarchical monitoring and giving residual revenue rights to the central monitor. This can be seen as an argument for providing equity compensation to management but not to employees more broadly.

Core and Guay (2001) argue that firms with severe cash constraints and high capital needs may substitute equity compensation for cash pay in order to ease liquidity constraints. For instance, ICT companies that have not yet secured positive income streams and are investing heavily relative to their assets may use equity-based pay for this reason.

Based on Finnish data, Pasternack (2002) suggests that if options are believed to solve the principal-agent problem between owners and managers, then the presence of significant foreign ownership would increase the probability of observing option schemes. An alternative explanation for a positive relationship is that, initially, foreign owners can be more familiar with such schemes than are others (Huolman et al., 2000).

Ittner, Lambert and Larcker (2003) emphasize that more concentrated equity ownership would decrease the likelihood of observing stock option schemes, since large shareholders can resort to alternative means of monitoring the management. The managerial power approach of Bebchuk and Fried (2003) predicts that since managers are more weakly monitored under dispersed ownership, they are more likely to grant themselves options. Concentrated share ownership may also reduce share liquidity, which inhibits information production in the stock market, distorting the signals from share prices and reducing the attractiveness of equity-based compensation measures such as stock options (Holmström and Tirole, 1993). Thus we may expect ownership concentration to be negatively related to the use of options.

The gain in share values may not be the only channel how equity prices influence the adoption of option schemes. Kalmi (2005) shows in a theoretical model that firms having higher levels of market values per employee are more likely to use stock options as a compensation method and are also more likely to include the overall workforce in the scheme. When market values fall and options become a less cost-efficient means of remuneration, options will be targeted to a more select group of the workforce.

This suggests that in a declining stock market, the number of option schemes in general will diminish, and that this is particularly the case with broad-based option schemes.

Murphy (2002) suggests that firms try to offset under-water option schemes by launching new schemes where exercise prices are considerably lower than in previous programs. This argument suggests that the stock market decline would not decrease the number of new option plans as firms update their schemes to reflect the new stock market realities.

In the first essay we contribute to this literature by testing all the adoption explanations described above. Our main focus is to examine how firm characteristics and stock market conditions influence adoption patterns with a desire to maintain comparability with the previous literature that mainly uses U.S. data. We use a new, long, and rich panel data set consisting of all Finnish publicly traded firms from 1992 to 2003 to provide evidence on a number of hypotheses relating to the adoption of selective versus broad-based stock option schemes. At the stock market level, we find that general patterns concerning the adoption of option schemes correlate strongly with overall market developments. Stock option adoption seems to be a procyclical phenomenon, becoming more common and inclusive during a stock market upturn, and less common and selective during a downturn. These findings are consistent both with an explanation that stresses changes in market values as a primary explanatory variable as well as an explanation based on the importance of the levels of market values. However, firm-level analysis enables us to probe deeper concerning the two competing explanations. In this firm-level analysis, we find that higher market values per employee lead to higher probabilities of the adoption of both broad-based schemes and selective schemes, while higher returns predicts exclusively selective schemes.

We also test extant hypotheses concerning the impact of firm characteristics on the adoption of stock options. Consistent with previous results in the literature, our findings suggest that the adoption of broad-based plans is related to difficulties in monitoring employee performance. Another finding is that larger firms as well as firms with more dispersed ownership are more likely to adopt selective schemes. This finding is consistent with the hypothesis that options schemes are adopted in situations where the agency problem between owners and management is especially severe. We also find that share returns from the past year affect the adoption of targeted stock options, but that there is no effect on broad-based plans. We find only weak evidence for a liquidity constraint hypothesis. Finally, we find no evidence for the hypothesis that foreign ownership influences the adoption of options schemes.

The second essay examines the productivity effects of broad-based and selective stock option schemes. One of the key arguments of the proponents of option schemes has been that options can align the interest of employees with those of shareholders. For example, options may motivate employees to exert more effort and take actions that are mutually beneficial to both owners and employees. The motivational effect may be especially relevant in situations where alternative approaches, such as direct monitoring or piece rates, are not feasible due to monitoring difficulties (Sesil, Kroumova, Blasi and Kruse, 2002). Stock options may also help to retain current key employees, both because they adjust pay according to the current market conditions (Oyer, 2004) and because they are a deferred form of compensation. In addition, options may help to recruit new employees.

Nevertheless, there has also been a strong criticism of stock options. For example, options, when exercised, entail a cost to shareholders in the form of dilution of ownership. Also others note that the costs of stock options have not been included in income statements (e.g. Hall and Murphy, 2003). Hence they argue that the increasing popularity of options in part reflects firms' mistaken thinking that options are a cheaper form of compensation than their true cost. Also, payment schemes that reward collective performance suffer from a free-rider problem: an individual who increases his effort will bear the full cost of the increase in effort, but will realize only a small part of the resulting increase in output (e.g. Alchian and Demsetz, 1972; Oyer, 2004). Another type of criticism comes from the psychological expectancy theory (Vroom, 1995): according to the "line-of-sight" argument, rewards based on performance can only be motivating if, by their actions, employees can influence the measures on which performance-pay is based. Since especially the effects of an individual non-managerial employee's actions on a firm's stock price can be expected to be marginal, according to this view the motivational impact of options is minimal. Although considerable theoretical research has been done on stock option schemes, ultimately the question of the productivity impacts of options is an empirical one.

When turning to empirical work, it appears that there is only a limited amount of work that examines the consequences of stock options for firm performance. Conyon and Freeman (2004) focus on the economic outcomes of broad-based option schemes in a sample of U.K. listed firms in 1995-98. They use three survey data sets for 284 firms and find evidence that the presence of a stock option plan is significantly associated with higher firm-level productivity.

Sesil, Kroumova, Blasi and Kruse (2000) use survey data on broad-based option schemes in 1997 from different sectors in the U.S. provided by the National Center for Employee Ownership. They find that in stock option firms productivity is 28% higher than non-option firms and 31% higher than their non-option pairs. However, the response rate of the survey is only 10% yielding 73 firms in the final data set.

In a subsequent study the same authors (Sesil, Kroumova, Blasi and Kruse, 2002) compare the performance of 229 new economy firms, which offer broad-based option schemes to their non-stock option granting counterparts. They find evidence that productivity is higher in firms with broad-based plans. Their research methodologies include descriptive analysis, paired matching comparisons between broad-based and non-broad-based option firms within the same industry, and a cross-section regression for 1997.

Ittner, Lambert and Larcker (2003) use survey data to examine 217 new economy firms during 1999-2000. By using cross-section analyses they examine the performance consequences of option and equity grants to senior-level executives, lower-level managers, and other employees. Their findings indicate that lower than expected option grants and/or existing option holdings are associated with lower accounting and stock price performance in subsequent years.

In sum, the empirical evidence from these studies suggests the existence of a positive and often quite sizeable link between broad-based stock option plans and productivity at the firm-level. This implies that the available empirical evidence provides support for theorists who predict that potentially powerful economic effects will flow from options and dominate the effect of factors such as free riding, accounting myopia and managerial rent-seeking. However, before accepting this conclusion it is important to note some key shortcomings of these studies. For one thing, in all of these studies, often the data are based on surveys and are apt to suffer from various selection biases, for example three studies focus only on new economy firms. Second, nearly all studies are limited insofar as they concentrate on the time-period before the stock market slump in 2000. Third, and at odds with most theories, no studies are able to reliably distinguish the productivity impact of selective versus broad-based plans. Finally, there is the issue of the appropriate econometric approach. Amongst several potential matters, the sensitivity of findings to potential issues of endogeneity of option schemes and inputs used in production is clear.

In the second essay we respond to these issues as best we can. Our data will enable us to address most of these matters. The only exception is our lacking data for other HR practices, such as employee participation in decision making. Because we utilise public firm-level data on stock option plans, we cannot control for the level of employee participation in decision making, the existence of profit sharing, and other human resource management practices. We can, however, separate firm fixed effects and common time-specific effects from several other factors that possibly have effects on firm productivity using the fixed effects and the GMM estimators.

The most important finding, yielded almost consistently in diverse specifications, is a statistically insignificant association between option programs and firm productivity. This result is exceptionally robust for broad-based schemes and is independent of what option program indicator is used in estimations. As such our findings are consistent with those who hypothesize that the performance impact of options may be limited because of reasons such as free-rider problems (e.g. Oyer, 2004), accounting myopia (e.g. Hall and Murphy, 2003) or line-of-sight arguments (e.g. Vroom, 1995). In addition, our results are consistent with much of the financial literature that does not find evidence of a link between options and business performance (e.g. Hall and Murphy, 2003.)

For selective programs, however, the findings are less consistent. In our baseline fixed effects estimates we find a statistically significant productivity impact that is between 2.1-2.4%. Since most selective plans are allocated to executives and/or key employees, this finding also provides support for the line-of-sight argument – rewards based on performance can only be motivating if the action of employees can influence the measures on which the performance-pay is based. Equally, this evidence does not support those who stress managerial rent-seeking as the principal reason for the adoption of a selective option plan. However, in models where endogeneity is accounted for, we do not find any strong evidence of a link with firm productivity. Similarly when we take into account the dynamics of selective programs, no association between selective options and firm productivity is found. In sum, our findings do not provide strong support for the hypothesis of a positive association between option schemes and firm productivity.

The third essay provides first empirical evidence on the relationship between stock option schemes and firm technical inefficiency in the stochastic frontier literature. Whereas sharp disagreements exist among theorists on the economic impact of different types of option schemes, an existing empirical work on the performance effects of options in economics has typically focused on the link between options and firm productivity. For example, we argue in Essay 2 that: *"For selective option schemes, the baseline fixed effects estimator suggests a 2.1-2.4% positive and statistically significant effect of the option program (size) indicator on firm productivity. However, in empirical models in which endogeneity and dynamics are taken into account, no evidence is found of a link with firm productivity."* As a consequence the evidence of a non-significant link with firm productivity raises a question whether, instead of productivity, there is a link between stock options and firm technical inefficiency. For example, one may expect that the adoption of an option scheme motivates a firm's managers and employees to make better decisions, work harder and share information within a firm in a way that reduces firm inefficiency. On the other hand, it can also be argued that options increase technical inefficiency, e.g. due to free riding, accounting myopia and managerial rent-seeking.

From an empirical point of view, a research question is how an exogenous factor, such as stock options, affects firm technical inefficiency. We use maximum likelihood stochastic production frontier estimators that provide the parameterizations of exogenous influences on the mean and the variance of firm technical inefficiency. The key research questions are the following: (i) whether firm-level inefficiency is higher in non-option than in option firms; (ii) whether the impact of options on firm inefficiency is dependent upon whether the plan is broad-based or selective. In the empirical analysis we use novel panel data of Finnish publicly listed firms in the manufacturing and the ICT sectors from 1992 to 2002. Our unique data enable a careful investigation of the inefficiency effects of different types of option plans.

We find that the shape of the inefficiency distribution differs notably between the manufacturing and ICT sectors. For example, the mean inefficiency estimates in the ICT sector are substantially higher than in the manufacturing sector, though naturally efficient and inefficient firms exist in both sectors. Also, in the ICT sector the conditional mean inefficiency estimates indicate that there is no inefficiency difference between option and non-option firms. However, in the manufacturing sector our findings suggest that broad-based firms may have higher mean inefficiency than selective and non-option firms.

The quantitative assessment of the average marginal effects on the inefficiency term supports the view that especially broad-based schemes may affect the mean and the variance of the inefficiency term in the manufacturing sector. The findings on the mean of inefficiency suggest that broad-based schemes may increase technical inefficiency. Respectively, the average marginal effect of broad-based schemes on the variance of the inefficiency term is significant implying an increase in production uncertainty. In sum, these findings would indicate, other things equal, that broad-based scheme firms in the manufacturing sector may achieve lower and more uncertain productivity growth as time goes by. For selective schemes, we find no evidence of a link with technical inefficiency. Finally, our findings do not support the hypothesis that option schemes may reduce firm technical inefficiency.

The fourth essay studies how CEO compensation is related to firm size and to firm performance. The previous empirical literature has emerged mainly during the last 25 years, since before the 1980s only a handful of academic studies of CEO compensation were published. The evolving literature has been truly interdisciplinary: an extensive number of CEO compensation studies have been conducted in economics, finance, accounting and management.

Jensen and Murphy (1990) suggest that CEO wealth changes $3.25 for every $1,000 change in shareholder wealth in the U.S. In addition, they argue that CEO pay-for-performance sensitivity has been modest and it has fallen in real terms from the 1930s: "*... on average, corporate America pays its most important leaders like bureaucrats ... The total change in all CEO wealth is $3.25 per $1,000 change in shareholder wealth for the full sample, $1.85 for large firms, and $8.05 for small firms. The largest CEO performance incentives come from ownership of their firm's stock.*"

Rosen (1990) surveys several empirical studies on CEO compensation. The evidence suggests that the effect of stock returns on log compensation is in the range of 0.10-0.15. He also summarizes a variety of pay-for-firm size elasticity studies for different time periods in the U.S. and the U.K. He finds some variation in CEO pay-for-firm size elasticities, but *"...the relative uniformity of estimates across firms, industries, countries, and periods of time is notable and puzzling because the technology that sustains control and scale should vary across these disparate units of comparison. The estimated elasticities for all companies are not significantly different from 0.3."*

Gregg, Machin and Szymanski (1993) focus on the relation between the wage of the highest paid director and firm performance with U.K. data for 288 large listed firms from 1983 to 1991. They find that the link between the top director's pay and firm performance is very weak in the terms of share returns. However, after splitting the sample into two sub-periods, i.e. 1983-1988 and 1989-1991 (recession period), they find a positive but small pay for performance link for the first sub-period, but not for the second. In addition, they argue that growth in the top director's pay is strongly correlated with firm growth: a 50% increase in sales leads to a 10% increase in compensation.

Conyon and Leech (1994) examine the determinants of the top director's salary and bonus with a sample of 294 large U.K. listed firms in 1983-86. They find a positive but very small pay elasticity estimate with respect to firm performance. For the median director, a 10% increase in shareholder wealth corresponds to an increase in compensation of 375 pounds. Another key finding is that both ownership control and concentration decrease the level of the top director's pay, but these variables do not affect the growth of pay.

Main, Bruce and Buck (1996) use panel data for 60 U.K. firms from 1981 to 1989. They find that because of stock options there is a statistically significant relationship between the wage of the highest paid executive and firm performance. For example, a 10% increase in shareholder wealth increases the top paid director's compensation about 9%. The sensitivity of top executive compensation with respect to firm performance is greater than in the previous U.K. studies, since they have also taken into account information on stock options.

Hall and Liebman (1998) use a 15-year panel data set of CEOs in the largest U.S. firms from 1980 to 1994. They argue that CEO compensation is highly responsive to firm performance if the value changes of CEO stock and option holdings are taken into account in the empirical analysis. For example, the median elasticity of CEO compensation with respect to firm market value is 3.9 for 1994, which is about 30 times larger than previous estimates that rely on salary and bonus changes alone. They also argue that CEO mean (median) compensation increased by 207% (146%) in real terms between 1980 and 1994. Perhaps more importantly, virtually all of this increase is attributable to changes in the value of CEO holdings of stock and stock options. When using an analogous measure to that of Jensen and Murphy (1990), in 1994 the total change in CEO wealth is $5.25 per $1,000 change in shareholder wealth. Although this degree of sensitivity may appear modest, Hall and Liebman show that CEO wealth may change millions of dollars for typical changes in firm value. Thus, they conclude that CEO compensation is strongly related to the success of the companies they manage.

Murphy (1999) provides support for the key role of stock options in his broad survey on CEO compensation studies: *"... our analysis shows that CEO pay-performance sensitivity has nearly doubled to $6.0 per $1,000 change in shareholder* *value by 1996. The increase in pay-performance sensitivities has been driven almost exclusively by stock option grants."*

Randøy and Nielsen (2002) examine the relationship between firm performance, corporate governance and CEO compensation in Sweden and Norway in 1998. Their findings indicate a positive relationship between the size of the board and CEO compensation, foreign board membership and CEO compensation, and firm market capitalization and CEO compensation. On the contrary, they find a statistically insignificant link between firm performance and CEO compensation.

Kato and Kubo (2005) examine the link between CEO compensation and firm performance in Japan with a panel data set from 1986 to 1995. They find that CEO cash compensation is sensitive to firm performance, especially on accounting-based measures of performance. On the contrary, stock market-based performance seems to be a less important factor in assessing CEO compensation sensitivity. One reason may be the fact that until 1997 executives' stock options were banned in Japan, except at small venture capital companies.

In Finland, Vittaniemi (1997) has studied the relationship between CEO compensation and firm performance previously. He uses panel data on 48 listed and 70 non-listed firms in 1989-93 (5 years) and estimates separate models for listed and non-listed firms. The key finding is a significant CEO pay-for-performance relationship in listed firms, based on once lagged performance measures. However, in non-listed firms the relationship is less important.

Besides CEO pay-for-performance, the previous empirical studies have focused on the link between CEO compensation and firm size. The evidence from the literature suggests a relative uniformity of CEO pay-for-firm size elasticity point estimates of 0.3

across firms, industries, countries, and periods of time. For example, Baker, Jensen and Murphy (1988) report cash compensation elasticities with respect to firm sales in the range of 0.25-0.35, when summarizing the U.S. Conference Board data from 1973 to 1983. Rosen (1990) supports this finding by summarizing a variety of studies for different time periods in the U.S. and the U.K. Also, the estimates of Conyon and Murphy (2000) with the U.K. and the U.S. data in 1997 support "the near uniformity elasticity hypothesis of 0.3" for the U.S., but not for the U.K. firms.

In the fourth essay we contribute to this literature using individual-level CEO compensation data from Finland between 1996 and 2002. By providing new evidence from a very different institutional context than the U.S. and the U.K., we hope to increase our understanding of CEO compensation practices across different countries. We follow previous empirical studies in the literature by exploring CEO pay-for-firm size elasticity and CEO pay-for-firm performance sensitivity. We estimate several empirical specifications, where we control for the industry of the firm, CEO age, the size of the board, the voting share of a largest shareholder and the share of foreign ownership, since all these variables may affect the level and the changes of CEO compensation.

Our key finding is that CEO average compensation has increased substantially between 1996 and 2002. For example, the ratio between CEO and industrial worker average total compensation was 7 in 1996, peaked at 24 in 2000, and thereafter dropped to 13 in 2002. In addition, CEO mean salary and bonus (in real terms) was about €166,000 (median €147,000) in 1996, whereas it was about €280,000 in 2002 (median €208,000). The percentage increase from 1996 to 2002 is 69% (median 41%). Respectively, CEO mean total compensation increased approximately from €180,000 to €357,000 (98%), whereas median total compensation increased approximately from €155,000 to €233,000 (50%) in the period.

According to our estimates, CEO average compensation, and especially average total compensation, is highly related to stock market-based measures of firm performance, such as shareholder wealth and share return. For example, shareholder wealth, close to that of Jensen and Murphy (1990), suggests that the salary and bonus change in CEO wealth is €6.84 per €1,000 change in shareholder wealth. Respectively, the total compensation change is €21.85 per €1,000 change in shareholder wealth, likely due to a few large stock option exercises. The estimated "semi-elasticity" of CEO salary and bonus with respect to share return is 0.09, and 0.28 for total compensation. We find no evidence on the contemporaneous link between the change in CEO compensation and change in ROA% (Return on Assets), an accounting-based measure of firm performance. However, changes in one year lagged performance measures, both accounting and stock market-based, can be associated with the change in CEO total compensation. In addition, CEO pay level is significantly related to firm size. Typically, pay-for-firm size elasticity is in the range of 0.2-0.3.

Finally, we also find some interesting corporate governance findings. First, the share of foreign ownership is positively associated with the level of CEO compensation. In most specifications, the parameter estimates of foreign ownership are about three times larger for total compensation than for salary and bonus. Second, ownership concentration, as measured by the voting share of a largest shareholder, is negatively related to the level of compensation in the pooled model. Thus this finding is consistent with hypotheses that highlight large shareholders' possibilities to monitor executives. Third, the size of the board is positively related to the level of compensation, especially to base salary and bonus. This supports the hypothesis that underlines inefficiency, rent-seeking and free-rider issues that can be associated with a sizable board.

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The Determinants of Stock Option Compensation: Evidence from Finland[1]

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Abstract:
A new, long and rich panel data set consisting of all Finnish publicly traded firms is used to study how firm characteristics and stock market developments influence the adoption and targeting of stock option compensation. Stock option adoption is found to be a pro-cyclical phenomenon. Findings from firm-level econometric analysis often corroborate those based on U.S. data, but important differences also emerge. Findings include: (i) firms with higher market value per employee are more likely to use stock option compensation; (ii) share returns from the past year affect the adoption of targeted stock options, but not broad-based plans; (iii) our results are consistent with the hypothesis that selective and broad-based plans arise as solutions to differing monitoring difficulties. Broad-based schemes are observed when production is human capital –intensive and employee performance is hard to monitor, while selective schemes are adopted when ownership is dispersed and therefore owners may have weak incentives to monitor the management.

Keywords: stock options, compensation, corporate governance

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1. Introduction

During the 1990s, stock options have become an increasingly popular compensation method in a wide range of countries (e.g. Hall, 1998; Murphy, 1999). Although stock options were initially associated mainly with managerial compensation, this changed rapidly after more and more companies worldwide started to issue stock options to the workforce more broadly (Blasi et al., 2003). In turn this growth of stock options has generated heated public discussion with some viewing stock options as a device by which managers transfer excessive benefits to themselves, while others see options as a major innovation in managerial and personnel compensation.

In this paper, we examine the adoption of selective and broad-based stock option programs. A main focus is to examine how firm characteristics and stock market conditions influence adoption patterns. In addressing these questions we assemble and then use an exceptionally rich, new and long panel data set. Whereas most of the published literature uses U.S. data (and generally has been forced to rely on non-representative samples) our data are for *all* publicly traded firms in the interesting case of Finland. Most importantly, since our data span the years 1992-2003, for the first time we are able to analyse how stock market downturns, as well as upturns, affect the popularity of option schemes.

Our rich data enables us to investigate a large number of firm-level hypotheses and to see if findings based mainly on evidence generated mainly using US data are applicable to another country with a very different institutional setting. In addition, we pay special attention to the links between market value of equity and stock option compensation. While some recent literature has stressed the role of *changes* in market value and stock option compensation, we argue that there might be also an overlooked link between the *level* of market value (per employee) and stock option compensation. Firms with higher level of market value per employee are able to provide incentives at lower cost, thus encouraging the use of stock option compensation. This has implications both at the level of the stock market and at the level of the firm that we address in our empirical research. Moreover, by providing new evidence for firms that exist in a very different institutional context than the U.S. we examine the generality of conclusions concerning the impact of various firm characteristics on the adoption of different forms of stock options.

2. Conceptual framework

In this paper, we examine the determinants of whether a firm opts for a *selective* or a *broad-based* option scheme. By selective schemes we mean schemes that are targeted to selected group(s) among the workforce. These include managerial schemes, but also schemes targeted to key personnel (e.g. R&D workers). In broad-based schemes, the majority of the workforce is eligible to participate. Broad-based schemes are all encompassing, including managers, and they do not have to be egalitarian in the sense of all participants receiving the same number of options.[2]

There is a rich literature on various issues related to stock options. Some of the earlier literature was concerned on the use of stock options in managerial compensation (e.g. Jensen and Murphy, 1990; Hall and Liebman, 1998; Conyon and Murphy, 2000; Mäkinen, 2005; see also the survey by Murphy, 1999). In finance, there have been studies concerning the timing of announcing and exercising stock options (e.g. Huddart and Lang, 1996; Yermack, 1997) and the valuation of employee stock options (Meulbrok, 2001; Hall and Murphy, 2002). After stock options diffused across wider range of personnel there has appeared studies relating to the performance effects of stock options (Sesil et al., 2002; Conyon and Freeman, 2004) as well as to the incidence and adoption of these plans (Core and Guay, 2001; Ittner et al., 2003; Kroumova and Sesil, forthcoming).[3] It is to the latter tradition of adoption and incidence our study connects to most closely.[4]

There is also another literature examining a closely related topic, namely the incidence of *employee stock ownership plans* (e.g. Jones and Kato, 1993; Kruse, 1996). In the following by carefully examining this broad body of work we identify a number of hypotheses that we will subsequently test with our new data set. Our discussion proceeds with an eye to subsequent model selection and a desire to maintain comparability with previous empirical research on stock options.[5]

Share return performance. Bergman and Jenter (2004) develop an elaborate argument to try to account for the puzzling observation that stock options are used at all to compensate risk-averse employees. They argue that it is necessary that employees have more optimistic expectations concerning stock price developments than do outside investors. They attribute this optimism to "excessive extrapolation", where employees form expectations based on past share returns and believe that high returns will continue in the future. Thus firms with high share returns find it cheaper to pay employees partly in equity instead of cash, and these equity payments are valued more by employees than by risk-neutral investors. The argument requires a degree of irrationality on the part of employees, compared to the rational expectation benchmark provided by outside investors.[6]

Market value per employee. The growth in share values may not be the only channel how equity prices influence the adoption of option schemes. It is also possible that firms with higher *levels* of share values are more likely to provide option compensation. To understand the basic idea, consider two firms that have identical growth prospects. However, in firm A the market value of equity per employee is 2000 units, while in firm B it is 1000 units (perhaps because the work is less physical capital–intensive than in firm A). The expected growth rate of market value of equity is 10 % for both firms. Giving employees options corresponding to 10 % of total equity would yield an expected pay-off of 20 units (2000*0.1*0.1) in firm A and 10 units in firm B. Thus, in firm A the value of a given amount of equity compensation is higher than in firm B or, alternatively, firm A can provide the same expected compensation to its employees as firm B with lower dilution costs to shareholders. Notice that it is the market value per employee, rather than market value *per se*, which is relevant for compensation purposes. Kalmi (2005) has shown in a formal model that firms that have higher levels of market values per employee are more likely to use stock options as a compensation method, and also more likely to include the overall workforce in the scheme.

Both the argument concerning the growth in market values and the argument on the level of market value per employee suggest that stock options would be used especially during bull markets. However, these two arguments have different implications concerning what the key variable is in explaining the adoption of broad-based stock options, so we can differentiate between the two models in firm-level econometric analysis.

To the best of our knowledge the only paper that discusses what happened to the stock option compensation when stock prices fell recently is Murphy (2002). He suggests that many option schemes went underwater when stock values declined and thus provided virtually no incentive effects. Consequently, firms tried to offset these declines by launching new option schemes where exercise prices were considerably lower than in previous programs. This argument suggests that the stock market decline would not decrease the number of new option plans, and perhaps would increase them temporarily, while firms would update their schemes to reflect the new stock market realities. However, the argument on market value per employee outlined above suggests that option schemes would become more expensive for shareholders when market values decline. When market values fall and options become a less cost-efficient means of remuneration, the model of Kalmi (2005) predicts that options will be targeted to a more select group of the workforce. This follows from the assumption that the marginal impact of stock options on incentives is higher for managers than for lower-level employees. This suggests that in a declining stock market, the number of option schemes in general will diminish, and that this is particularly the case with broad-based option schemes.[7]

Risk. A standard prediction from the principal-agent theory suggests that, *ceteris paribus,* risk-averse employees would dislike schemes where a part of pay is tied to a volatile measure, and where a significant part of the volatility is beyond the control of employees. However, Prendergast (2002) has argued that empirical research often finds a positive relation between risk and incentives, since in more uncertain settings the principal is often better off delegating responsibility to the agent(s), and the delegation necessitates the use of incentives. Oyer (2004) argues that when uncertainty is high, fixed wage contracts require frequent revision, but the transaction costs of rewriting the contracts become prohibitively costly. To retain the best employees, it is better to tie compensation to a measure that correlates with the business cycle. Again this gives rise to a positive correlation between risk and option compensation, which Oyer and Schaefer (2005) observe in their empirical analysis. In sum, theoretical predictions concerning the relationship between risk and options use appear to be rather ambiguous.

The effect of group size. The famous 1/n problem (Alchian and Demsetz, 1972) suggests that group-based incentive schemes, such as equity pay, become ineffective when the size of the group grows. According to Alchian and Demsetz, larger groups necessitate hierarchical monitoring and giving residual revenue rights to the central monitor. This can be seen as an argument for providing equity compensation to management but not to employees more broadly. This solution assumes that managers can monitor employees at low cost, but the principal-agent problem between owners and management requires the use of performance incentives. However, empirical research on share schemes and profit sharing has typically found that the likelihood of sharing schemes increases with firm size, rather than vice versa (e.g. Jones et al., 1997; Sesil et al., 2003). Possible explanations for this anomaly are either the existence of fixed costs associated with establishing the schemes, or mutual monitoring of employees or employee co-operation that eliminates the negative impact from firm size (e.g. Blasi et al., 1996).

Human capital intensity and monitoring of employee performance. Earlier research suggests that human capital intensity in the production process should be an important determinant of option plans (Core and Guay, 2001; Kroumova and Sesil, forthcoming). When work is human capital intensive, it becomes difficult to monitor and, in lieu of direct supervision, employees need self-motivating incentive schemes. Such situation arises when the level of intangible assets is high.

Liquidity constraints. According to Core and Guay (2001), firms with severe cash constraints and high capital needs may substitute equity compensation for cash pay. For instance IT companies that have not yet secured positive income streams and are investing heavily relative to their assets may use equity based pay for this reason.

Foreign ownership. If options are believed to solve the principal-agent problem between owners and managers, then the presence of significant foreign ownership would increase the probability of observing option schemes (Pasternack, 2002). An alternative explanation as to why we might expect a positive relationship is that, initially, foreign owners are more familiar with such schemes than are others, since such schemes were often imported from the U.S. (Huolman et al., 2000.)

Ownership concentration. According to the principal-agent theory, more concentrated equity ownership would decrease the likelihood of stock options, since large shareholders can resort to alternative means of monitoring the management (Ittner et al., 2003). However, when ownership is highly dispersed, each owner has only weak incentives to monitor the management, therefore increasing the need for self-enforcing incentive schemes.[8] Concentrated share ownership may also reduce share liquidity which inhibits information production in the stock market, distorting the signals from share prices and reducing the attractiveness of equity-based compensation measures such as stock options (Holmström and Tirole, 1993). Thus we expect ownership concentration to be negatively related to the use of options.

3. Institutional environment

In this section, we briefly review relevant institutional developments, paying special attention to the development of corporate governance, industrial relations, and taxation. In the main we find that developments in corporate governance and industrial relations have been favourable to the adoption of option schemes. However, the effects from taxation have been detrimental, though apparently not of decisive importance.

In the end of 1980s, the Finnish corporate governance system in listed firms was very much bank-centred and resembled the German system.[9] Financial institutions owned around 25% of the value of shares in the Finnish stock exchange. Bank loans were the most significant source of external funding for listed companies. On the other hand, at the end of the 1980s, the stock market was booming and the number of firms listing was at a record high.

During the early 1990s Finland suffered the most severe depression in any OECD country since World War II. For example, during 1990-1993 unemployment soared close to 20 % and GDP plummeted by 14 % (Kiander and Vartia, 1996). This also caused a significant change in financial markets: bank loans dropped significantly, as did share prices. After the devaluation in 1991 and the floating of Finnish currency Markka in 1992, the stock market started its recovery, but bank lending continued to decline throughout the 1990s. Nowadays, the equity market has become an important source for external funding for publicly traded firms, and Finland has shifted from a bank-based financial intermediation to a market-based system.

Turnover on the Helsinki Stock Exchange grew dramatically during the 1990s (although this is partly because of the growth of Nokia) and the number of firms listed also increased significantly, especially in the late 1990s and 2000. No doubt this has contributed to the prevalence of option schemes. Now stock markets are much thicker, more informative and more transparent. This reduces the possibility that managers may manipulate stock prices and that options would only be an instrument in self-serving deals by managers. At the same time, both monitoring of insider trading and legal punishments have become stricter.

Another important development is the increase of foreign ownership. The Finnish stock market was opened to foreign investors only in 1992, but today foreigners are the largest ownership group (although this is largely because of Nokia). By 2000, foreign ownership had increased to 53%, while ownership by domestic financial institutions had dropped at the same time from 20% to 4% (Hyytinen et al., 2003). The increase in foreign ownership has contributed to the transformation of the Finnish business towards a more competitive and open culture where shareholder value is given a high priority (Tainio and Lilja, 2003). As noted above, foreign owners may have also played a large role in demanding that firms use stock options.

Finally, we note that the largest increase in the use of options took place in 1998-2000, when the stock market was at record highs. Table 1 depicts the growth in share prices at the Helsinki Stock Exchange between 1990 and 2002. The difference between the "general index" and the "portfolio index" is that in the latter the maximum weight of any one company is limited to 10 % (the portfolio index is available only starting from 1996). After the decline in stock prices during the early 1990s, the stock market started its recovery in 1992, while 1993 was a year of extraordinarily good performance with stock prices almost doubling. After two more moderate years, extraordinary growth resumed in 1996. In 1999 the general index grew by a spectacular 167%. The portfolio index, where the impact of Nokia has been curtailed,[10] has behaved more moderately, but even the portfolio index rose over 72% in 1999. During the boom years many investors believed that, thanks to the arrival of the "new economy", stock prices would continuously rise. Investors' "over-optimism" may in part explain the increased use of options and also why owners were not very responsive to the concerns of shareholder value dilution. However, stock market prices started to fall after May 2000, accelerating further in 2001 and 2002.

[Table 1 about here]

Turning to industrial relations, we observe both continuity and change.[11] Consensual collective bargaining and centralised income agreements have continued as the norm for decades. Since the late 1960s, the unionisation rate of the workforce has been around 70-80%, and collective agreements are typically binding also for non-union workers or workplaces. Wage increases consist of a collectively agreed element that typically is economy-wide. In addition, firms can adapt their internal wage structures according to their financial possibilities. Throughout the 1990s, profit-sharing and other forms of performance-related pay have become common compensation methods throughout the economy (Kauhanen and Piekkola, 2002). Forms of performance-related pay are not negotiated in collective bargaining rounds, but employers can decide on their use unilaterally. The widespread use of performance-related pay, as well as the popularity of stock options, represents a change in industrial relations.

There is no obvious reason why the institutions of collective bargaining should affect the use of stock option schemes. The unions negotiate only on base pay, and they are not allowed to negotiate on option compensation. Thus, there is no direct influence from collective bargaining institutions. However, it is possible that by increasing employee bargaining power, employees may be better able to press the employers for the adoption of broad-based schemes. The strong role of the unions has also contributed to the relatively small wage dispersion, perhaps also favouring broad-based stock options. However, unions have very negative attitudes on stock options, which may also have reduced their use. In the late 1990's public opinion largely condemned managerial options while supporting the broadening of option schemes to the workforce at large.[12]

While the impact of changes in corporate governance and the transformation of industrial relations have benefited the diffusion of stock options, changes in tax laws and regulations have not been supportive. Of special importance are the changes that occurred in 1995; until then options were taxed as capital gains, thereafter they have been taxed as income. Typically this means (at least for all managers and most employees) that the tax rate is in the highest possible bracket (namely 58 %), compared to a top capital gains tax rate of 29% (recently dropped to 28%). Thus a significant tax disincentive to issue options was created, although in the end these adverse changes in taxation did not undermine the popularity of stock option compensation.

4. The increase in stock options in Finland

To provide comprehensive quantitative information on the nature and scope of stock option plans for all publicly traded firms in Finland, we integrate data from several sources. The initial attempt to organize the option data was by Professor Seppo Ikäheimo.[13] Since the original data were only up to 1998, we (and others working with the same data) have subsequently complemented these initial data by drawing on annual reports and stock exchange reports. We checked the veracity of the resulting data in several ways including contacting several companies directly and by working with Alexander Corporate Finance, a company that has enormous expertise in setting up option programs in Finland. Thus we believe that our data are comprehensive. While there is a small possibility of unrecorded data, our strong sense is that this would occur at most in a handful of cases.

Financial data are obtained from a database maintained by Balance Consulting, while the data on foreign ownership and market values of companies are from the Helsinki Stock Exchange (HEX).[14] Ownership data are collected from Pörssitieto-handbooks[15] and from the annual reports of companies. Finally, the data of stock returns and volatility are provided by the department of finance and accounting, Helsinki School of Economics (originally from the HEX).[16]

While our discussion focuses on publicly traded firms, we believe that the omission of privately held firms is not an important limitation of our analysis, because the use of option schemes has been concentrated mainly in public firms.[17] This is a natural consequence of the fact that options can work properly only in situations where the value of shares can be assessed by the stock market. However, options are also used by firms that expect to get listed in the relatively near future, especially in firms within the ICT sector.

In the analysis that follows, we distinguish between broad-based schemes and selective schemes. The latter are mostly managerial schemes, although they can also include other (key) personnel. However, in order to qualify as a broad-based scheme, all employees (or at least a great majority of them) should be eligible.[18] The classification is based on the public stock exchange reports. Typically the firms report whether the stock option programme is targeted to managers only, to managers and selected group of key personnel, or to the workforce at large. This is based on the Finnish Law on Joint-Stock Companies which requires that the firms should report all the relevant conditions on the stock option schemes to the shareholders prior to the adoption of the stock option scheme (Law on Joint-Stock Companies, Ch. 4, §12b). The adoption of the stock option scheme always requires the approval of shareholder meeting. Giving misleading information on the conditions of the stock option scheme, such as eligibility, may lead to legal sanctions. Employees can easily monitor if the announced eligibility rules match the actual practices. If the employees were to detect that the firm misreports the eligibility rules, the trade unions were bound to take an action (and virtually all Finnish workplaces, smallest firms excluding, have a union representative). Finally, since shareholders have to accept new option plans in the shareholder meeting, misleading information on eligibility rules would generate loud protests in the shareholder meetings. Thus, we have every reason to believe that the eligibility rules are reported accurately.

High rates of eligibility do not yet automatically translate into high rates of participation, and firms are not required to report the participation ratios. However, we have several pieces of evidence supporting the contention that high eligibility lead to relatively high participation rates, and what is crucial for our arguments, broad-based eligibility leads into substantially higher participation than narrow eligibility. There is naturally some variation in participation rates that is partly related to whether employees actually have to make an investment in order to subscribe options. In some cases employees are required to give a zero-interest loan to the company when they subscribe to options. Then the company pays back the loan at face value after a certain period (e.g. 1-3 years). Thus, employees face a cost in terms of foregone interest and liquidity. However, this cost is typically far below any real value of the options. Moreover, not all companies that use this procedure, but essentially give options to employees for free.[19]

There is also some, although not comprehensive, evidence on the actual participation rates. University of Nijmegen in the Netherlands collected data under the commission of the DG V of the European Commission on financial participation in different European countries, including Finland.[20] In this dataset, there are 19 firms that reported the participation rate of their stock option scheme by 2001.[21] By comparing these data to our dataset, we find that we have fourteen of these firms have been classified as operating broad-based schemes at the given moment, while five firms operate only selective firms. Reassuringly, the mean participation rate in the Nijmegen data is 69 % for those firms classified in our data as having a broad-based scheme, while it is 13 % for those companies that are classified as selective schemes.[22] We also checked from annual reports whether the companies with broad-based stock option programs report similar participation rates when reporting publicly (and then being under legal obligation to provide accurate information). While we could not find many firms that provided information on participation, those we found were consistent with the survey results.[23] Finally, an expert we interviewed confirmed that there are dramatic differences in the participation rates of stock option schemes, depending on the eligibility.[24]

Further, indirect evidence from this issue comes from the data we have on the size of the plants, measured by potential dilution[25]. The average broad-based plan involves 5.2% of outstanding shares and new shares potentially available through options, while the size of the average narrow-based plan is 3.1% of outstanding and new shares. The difference in means is significant at the 1% level. This suggest not only that the participation rates are significantly higher in the broad-based schemes, but also that the shareholders face a higher potential cost when more employees are eligible to participate.[26]

The remainder of this section describes the evolution of option schemes among firms listed in the HEX. This is done to provide the reader with a better understanding of the development and prevalence of option schemes in Finland. Our discussion also includes reporting evidence for a simple test of the hypothesised correlation between stock market movements and options use.

Table 2 describes the evolution of stock option plans in publicly traded firms in Finland between 1987 (when the first personnel stock option scheme was launched) and 2003. Our data consists of firms that are traded on the Helsinki Stock Exchange.[27] Since 1997 HEX has taken over the smaller lists and also has started to operate two additional lists besides its main list: the "I" (Investor) -list and the "NM" (New Market)-list. The "I-list" consists of firms that are traded infrequently and are often majority-owned by large investors. The "NM" list consists of smaller IT and high technology firms, similar to the NASDAQ or the Neuer Markt in Frankfurt. Thus, we have information on the presence of option schemes on the main list throughout the period and on the minor lists since 1997. We do not have information on firms that have not been listed in the HEX. However, we have included option schemes prior to the listing for such firms that enter the HEX before 2002. Typically, these option programs are adopted close to the listing (often one or two years before the listing).

[Table 2 about here]

Column 1 in Table 2 gives the number of firms in the main list during the period of our focus, while column 2 gives the total number of observations including the two minor lists (from 1997). As is apparent from the Columns 1 and 2, the number of firms at HEX fluctuates a lot with the business cycle.

Column 3 indicates how many firms adopted their *first* option scheme in a given year. Column 4 indicates that 59 of 127 firms initially adopted broad-based schemes. Note that the first broad-based scheme (column 4) is not necessarily the first option scheme shown in column 3: in fact, in 47 of 59 cases the first broad-based scheme is also the first scheme in general. This means that in 80 cases, the first scheme has been a selective scheme, and that only 12 of 80 (or 15%) of firms that have first adopted a selective scheme have decided to broaden it later to the entire workforce. In other words, it appears that if a firm is to give stock options to a broad range of personnel, it is more likely to do this from the inception of an options scheme rather than extend a scheme that initially was a narrow plan.

From Column 3 we observe that while seven pioneering firms installed their stock option plans as early as the 1980s, very few plans were launched during the depression years of 1990-1993. The renewed interest in option plans began in 1994 when 20 firms (almost one third of listed firms at that time) adopted option schemes. Relatively few firms adopted schemes during 1995-1996 (possibly because of the adverse taxation changes described earlier), but since 1997 options became widely popular again. The rise of option schemes during 1999-2000 is fuelled by new listings and when new listings stop after 2000, so does the introduction of new option schemes.

The first adoption of broad-based schemes has slightly different dynamics (Column 4). Although they have been used since 1989, they become popular only in

1998, when 14 firms adopted broad-based schemes. They retained their popularity until 2000.

Firms often launch new option schemes when existing schemes are due to expire, or they may operate many schemes simultaneously.[28] 86 of the 127 firms (68%) that had adopted an option scheme have installed more than one scheme (three firms have reached 7 successive schemes). An interesting finding about the successive option schemes (not apparent from the table) is that firms that initially choose a selective scheme are very likely to stick with that scheme. Of the 80 firms that installed a selective plan as their first plan, 55 (69%) had at least one successive scheme, and in 43 cases out of 55 (78%) all the successive schemes were targeted to a select group of personnel. In contrast, 31 of 47 (66%) firms that have a broad-based plan as their first plan installed at least one successive plan, but only in 7 cases of 31 (23%) were all the successive plans broad-based. Thus, while firms that first install a selective plan do not tend to broaden their plans, firms that initially adopt a broad-based stock option plan often subsequently adopt more selective schemes.

Column 5 shows the number of option schemes launched each year (including successive schemes) for firms in the main list and reveals dynamics that are very similar to the patterns reported in column 3. The early peak years are 1994 (16 main list firms adopt) and 1997 (17). In later years the adoption rate increases; in 1998 and 2000 more than 40 % of the main list firms adopted option schemes. However, after 2000 the adoption rate slows down with 25 new schemes in 2001, 22 in 2002, and only 15 in 2003.

Column 6 shows the number of broad-based schemes in the main list. The introduction of such schemes is concentrated in the years 1998-2000. In 1999 and 2000, almost half of the new option schemes in the main list are broad-based. However, when the stock market performance plunges, the adoption of broad-based schemes declines faster than that for selective schemes and only a handful of broad-based schemes are adopted in the main list after 2000. In total we identify 240 adoptions for main list firms between 1987 and 2003, of which 63 (26%) are broad-based.

Columns 7 and 8 use similar information as reported in columns 5 and 6, but do so for all firms. This data thus includes observations from minor lists, as well as observations from the pre-listing period for some main list firms. A total of 318 option adoptions are identified, of which 104 (33%) are broad-based. Interestingly, while 40 % of broad-based adoptions happen outside the main list, only slightly more than 10 % of selective schemes are adopted outside the main list.

In columns 9-12 we approach this issue from another angle and provide time series data on the existence of option schemes in main list firms (columns 9-10) and including all firms (columns 11-12). In these columns, we also use information on the timing of the scheme as well as on the launching of the scheme.[29] The data in column 9 indicate that the proportion of firms with an existing option scheme increases slowly but steadily until 1993, by which time a fifth of listed firms had an option scheme. This proportion jumps to around 40 % in 1994, after which it increases slowly for three years, until it starts to jumps again in 1998 to 65%. The temporary maximum is reached in 2001 when almost 85 % of the main list firms have an existing option scheme. Thereafter the proportion declines to 78 % in 2003.[30] Also the number of main list firms with broad-based schemes increases rapidly during 1998-2000, and in 2001 37 % of main list firms have broad-based schemes. This proportion declines to 34 % in 2003 (see column 10). Finally, columns 11-12 show developments for all firms, and also for those outside the main list.[31]

[Figure 1 about here]

In section II, we argued that the use of stock option compensation is likely to be related to the market value of firms. In Figure 1 we plot the stock price index (portfolio index) against the adoption new stock option plans and new broad-based option plans in the main list. The figure appears to provide support for the contention that increases in the stock market index and option use are related, although for broad-based plans this connection is not apparent before 1998 (due to infrequent use of broad-based plans before 1998.) We can examine this relationship more carefully by investigating the correlations between stock option use and market conditions. In the correlation analysis below, we restrict the analysis to main list firms, because we have comparable data only for main list firms. We calculate the correlation coefficient between the stock market portfolio index (general index used up to 1995), lagged by one year[32], and new stock options launched in the main list (as a percentage of all main list firms). The correlation coefficient is considerably large, 0.55. The correlation coefficient between the stock market performance and new broad-based stock options is also remarkably large, though slightly smaller, at 0.43. Broad-based stock options appear to be less sensitive to stock market performance because there were only few broad-based stock options before 1998. But after this point there has been a remarkable correlation between stock market conditions and the incidence of broad-based stock option plans. In general, there appears to be a large and significant correlation between the stock market conditions and the use of stock option compensation. In particular, our results show that there was indeed a significant drop in the number of option schemes after 2000, and that remaining schemes were targeted to a selective group of employees. This finding is entirely consistent with the hypothesis that the decline in market values would cause a decline in options issued and especially in the number of broad-based plans, but it is inconsistent with Murphy's (2002) previous empirical finding that during downturn, firms would replace their old under-water schemes with new schemes with lower exercise prices. Based on the results at the level of the stock exchange, it is not possible to determine whether the decline is due to the fall in market values or in expectations of share price development. We address this issue in the firm-level econometric analysis.

The use of stock option compensation in Finland appears to be comparable with the U.S. and other EU countries. Oyer and Schaefer (2005) cite figures from the B.L.S. survey on stock option grants in 1999. These statistics indicate that 12% of U.S. publicly traded firms issued stock options broadly to their employees in that year, and that broad-based schemes comprised 54% of all schemes in publicly traded firms. These figures are very comparable to our figures for Finnish main list firms in 1999 (15 % and 49%). For Europe, data for 2001 (Kalmi et al., 2004) show that the use of option schemes in countries including the UK, Germany, France and the Netherlands is comparable and in some cases even exceeds levels observed for Finland. However, apart from Finland, we do not know of any representative data showing how major movements in stock markets have influenced the use of stock options.

5. Econometric analysis

In testing hypotheses on the role of firm characteristics and stock market conditions in affecting the incidence of options, in the main we closely follow the approaches adopted in earlier work. This enables us to make comparisons with findings from previous studies. Thus we follow the literature concerning the way key variables are measured.[33] All monetary variables have been deflated by the consumer price index and, to address simultaneity concerns, all explanatory variables are measured for the year prior to the adoption decision.

Our data are an unbalanced panel of 799 observations for publicly traded firms at HEX.[34] The data are from the years 1992-2003. The data is unbalanced because many firms enter the stock market during the period of observations, and some exit the stock market. The number of firm-observations ranges from 18 in 1992 to 121 in 2000, with an average of 73 firm-observations by year. In total there are 127 firm observations.

Our data are exceptional in representing *all* the firms in the stock exchange and during a long time period (1992-2003). Many U.S. studies have relied on non-representative surveys and their data are for shorter time periods (e.g. Kroumova and Sesil, forthcoming; Ittner et al., 2003). While Execucomp data (e.g. Core and Guay, 2001) has the advantage of comprising data on individuals, since the data are restricted to managers, this means that aggregation to the firm level is not possible without guesswork. By contrast, our data are at the firm-level. The number of firm-observations is admittedly smaller than in some of the U.S. studies, but this is a constraint imposed by the size of the Helsinki Stock Exchange.

The basic econometric approach is to estimate multinomial logit models.[35] Our investigation focuses on the decision to *launch a scheme*. We believe this is the natural question that flows from our conceptual framework since *adoption* is likely to be very reflective of changes in stock market conditions and firm characteristics, unlike the *incidence* of a plan that is often fixed for years ahead. Furthermore, our key focus is on whether the firm decides to target its stock options either to a selected group or more broadly. Thus the dependent variable has three levels in our econometric models: 0 = firm i does not adopt any scheme at year t+1, 1= firm i adopts a selective scheme at year t+1, and 2 = firm i adopts a broad-based scheme at year t+1. In total there are 193 stock option scheme adoptions in this data set, of which 56 (29%) are broad-based schemes.

Our firm-level measures of independent variables are as follows:

Market value per employee is related to the cost efficiency of incentive provision, as explained above. It is included in logarithmic form.

Share return performance is measured by the continuously compounded daily stock market returns over a one-year period, in logarithmic form. This measure includes also dividends.

Risk is measured as the volatility (standard deviation) of daily stock market returns over a one-year period.

The effect of group size is proxied by the number of employees, measured in logarithmic form.

Human capital intensity. This is measured by using the ratio of intangible assets to the sum of tangible and intangible assets (as given on balance sheets). Thus a high level of intangible assets indicates that production is relatively human capital intensive, which in turn makes the monitoring of employee performance more difficult.[36]

Liquidity constraints. We use the interest rate burden (the ratio of net interest expenses to sales) as a measure of liquidity constraints. As a robustness check, we also have checked our results by using alternative measured suggested in the literature, such as cash level per employee and cash flow per assets. The results are not significantly altered by the choice of the proxy variable.

Ownership concentration is measured as the sum of the voting rights of the three largest owners (% of total votes). This variable addresses corporate governance concerns and the difficulties owners have in monitoring and disciplining management.

Foreign ownership is measured as the percentage of shares held by all foreign owners in that firm. Our data consists only of firms that are registered in Finland-- we do not include any subsidiaries of foreign firms. This variable also addresses corporate governance concerns.

Existence of previous option scheme takes the value 1 if the firm has an ongoing previous option scheme at year *t* and the value 0 if not.

The mean, standard deviation, minimum and maximum values for these variables appear in Table 3.

[Table 3 about here]

We proceed by piecewise augmentation of a basic specification. First, reflecting our focus on stock market conditions, we estimate a "baseline model" that includes only firm market value per employee, share returns, and volatility, as well as controls for industries and years. In the second stage we augment the baseline model and include several other variables that are identified in the literature and which mainly reflect other firm-level characteristics. Of key interest is to determine whether the results from the basic model survive after the inclusion of these additional variables. We assume that observations coming from the same firm are not independent, and adjust the standard errors for clustering within firm-observations.

We use the multinomial logit model where we estimate the probabilities for each three possible choices in the given year: selective scheme, broad-based scheme, or no scheme. The estimation results from the model are presented in Table 4 where two sets of parameter estimates corresponding to each specification are reported. The first set (A1, B1 etc.) indicates how much the probability that a firm adopts a selective scheme changes when the explanatory variable increases by one standard deviation. The second set (A2, B2 etc) of coefficients indicates a similar probability effect for the decision to adopt a broad-based scheme.

[Table 4 about here]

Looking at the reported coefficients in columns A1 and A2, the estimated baseline probability of adopting a selective scheme is 16.7% and the corresponding probability for a broad-based scheme is 5.2%. The variables are jointly significant. Market value per employee is strongly statistically significant for both the broad-based decision and also for selective schemes. A one-standard deviation increase in (the log of) market value per employee would increase the probability of observing a selective scheme by 4.9 percentage points, and the probability of observing a broad-based scheme by 4.1 percentage points. When the effect is translated into percentages (instead of percentage points), it is in fact much larger for broad-based schemes. Specifically, when (the log of) market value per employee increases by one standard deviation, the probability of adopting a selective scheme increases by 30% while the probability of observing a broad-based scheme increases by 80%. This pattern is consistent with the hypothesis that the costs of stock option compensation are lower for firms with higher market value per employee.

For selective schemes, annual share returns seem to matter also a great deal: a one-standard deviation change in annual returns increases the probability of observing a selective scheme by 4.5 percentage points, while there is no significant effect for broad-based schemes. Thus this result is contrary to the findings of Bergman and Jenter (2004), who suggest that the impact from share returns should be higher for broad-based schemes. Finally, a one standard deviation increase in annual volatility increases the probability of observing a broad-based adoption by 1.6 percentage points, while volatility has no effect on the likelihood of observing an adoption of a selective scheme. While the first result is at odds with the hypothesis that high risk is expected to decrease the use of volatile compensation for risk-averse employees, the result is not surprising since it has often been observed in literature.

In columns B1 and B2, we add two industry controls: a dummy for firms in information technology, telecommunications, and electronics and a dummy for other manufacturing firms (service is the industry reference group.) The main results remain unaffected. For broad-based schemes, the impact of market value per employee remains highly significant but it is now 2.6 percentage points and thus somewhat smaller than previously. The effect of market value per employee for selective schemes is estimated to be 4.4 percentage points. The impact from share returns on the likelihood of adopting a selective scheme increases somewhat to 5.0 percentage points, whereas the impact on broad-based schemes remains insignificant.

The next step of model specification is to include year dummies in the baseline model. These estimates are reported in columns C1 and C2. Reassuringly the addition of time dummies does not produce any marked changes in the reported findings.

Finally, in columns D1 and D2 we augment the baseline specifications reported in Table 4 with seven additional controls. The additional variables in this extended model are: number of employees (to control for the effects of group size); percentage of intangibles (to control for human capital intensity and resulting difficulties in monitoring employee performance); real interest burden (to control for liquidity constraints); to address corporate governance concerns, two measures of ownership, namely foreign ownership and ownership concentration; and a dummy for the presence of an earlier option scheme.

In the specifications reported in columns D1 and D2, the market value per employee variable remains significant, although the inclusion of additional variables causes the coefficient for broad-based schemes to fall to 1.5 percentage points. It is now significant at the 10% level. In contrast, for selective schemes the coefficient for market value remains approximately at the same level as before, at 4.4 percentage points, and it remains significant at the 5 % level.

The impact from share returns remains high for selective schemes but, as noted previously, share returns have no impact concerning the adoption of broad-based schemes. This is in sharp contrast with earlier research that has found that high returns lead firms to adopt broad-based plans. While earlier researchers have not included the market value per employee variable in their estimates, even if we drop the market value per employee variable from our regressions, we do not find a significant coefficient for the share returns variable in the broad-based specification. The finding that high returns predict selective schemes is somewhat of a puzzle, since it is at odds with previous findings. If we accept the argument that agents' expectations of future returns depend on current returns then, consistent with the argument of managerial opportunism presented by Bebchuk and Fried (2003), it may be that managers target options to a more select group of people when prospects are good.

By comparing the determinants of selective and broad-based schemes, it appears that both types of schemes reflect monitoring problems, but different ones. The coefficient for employment is positive and significant for selective schemes, while insignificant for broad-based schemes. This finding is consistent with the argument of Alchian and Demsetz (1972) that when group size is large, it is most effective to share the residual revenue rights only with management. Selective schemes are also used in large firms with widely dispersed ownership. The coefficient for ownership concentration is significant at 10% level and negative concerning selective schemes. This finding is consistent with the view that selective schemes are chosen in situations where the agency problem between management and owners is acute.[37] In contrast, the adoption of broad-based schemes is related to human capital intensity. This finding suggests that broad-based schemes can be found in situations where there is a high level of intangibles to all assets, indicating difficulties in monitoring employee performance.[38]

There are also some more unexpected and non-significant results. The coefficient for interest burden (a proxy for liquidity constraints) is not significant for broad-based schemes, but surprisingly it is significant for selective schemes (though only at 10% level). This finding would thus suggest that firms with liquidity constraints would use stock option compensation in order to ease their liquidity, but this finding does not carry over to broad-based schemes. While a possible interpretation, due to the fact that the result is neither very strong nor consistent with the prior expectations, it must be treated with some caution.[39] Volatility as a proxy for risk did have a significant coefficient in the first and the third specification, but this finding did not survive the inclusion of additional variables in the final specification. Given the ambiguous nature of theoretical predictions, this was not too surprising. Since previous work (Huolman et al, 2000; Pasternack, 2002) has identified foreign ownership as a major determinant of the adoption of Finnish option plans, surprisingly this variable was not found to be statistically significant in any of our specifications. However, since previous results apply to early experience with option schemes, it may be that the impact of foreign ownership on option adoption has declined in importance more recently. The dummy for a previous option scheme was insignificant. This may be due to conflicting influences: on the one hand not having an option scheme may indicate that a firm does not consider options to be a part of the optimal compensation strategy. In this case, the dummy should be positively related to the likelihood of the stock option adoption. On the other hand, having an option scheme in place may also reduce the need to adopt a new scheme, thus presenting a countervailing negative effect. Finally, the ICT and electronic sector dummy is not significant in any of the regressions, contrary to expectations. This may be due to multicollinearity with other variables in the specifications.

Since one company (Nokia) has contributed more than 50 % of the market value of the Helsinki Stock Exchange during most of the observation period, one might expect that the results would be sensitive to the inclusion of Nokia. Therefore we also estimate the multinomial logit specifications but without Nokia. To save space, we do not report those results here in detail.[40] However, the results are very similar to those reported in the paper and, indeed, some results are even more supportive of our hypotheses. For instance, market value per employee is statistically significant in every specification, at least at the 5 % level, for broad-based schemes.

Since share returns form a part of firm's market value, there is a possibility that the significant impact of market value per employee for selective schemes is partly driven by multicollinearity problems between that variable and share returns. To test this possibility, we replaced the current value of the market value per employee by its lagged value (measured at *t*-1). This is a good proxy for the current market value since the correlation coefficient between the two variables is high (0.84) and it is not correlated to the share returns. However, the downside is that since the information on lagged market value is missing for 106 observations, the sample size shrinks to 693. The number of broad-based stock adoptions decreases from 56 to 38. Keeping this in mind, when we re-estimated the specifications reported in Table 4, if anything these results are even more supportive of our basic hypotheses than those reported above. For example, while the results concerning share returns remained essentially similar and the impact of market value per employee remained similar concerning broad-based schemes, the coefficient for selective scheme decreased considerably and was no longer significant in the full model. In these results, the impact of lower cost of equity schemes leads to the adoption of broad-based, rather than selective, schemes. However, a robustness check of these results (i.e. with the current value of market value to employees replacing the lagged value but using the restricted, rather than the full, sample), finds that the results are essentially the same in the two sets of estimation using the restricted sample. In other words, the different results appear to be driven mainly by differences in the sample compositions, and do not mainly reflect the multicollinearity problem. This being the case, although we cannot completely exclude the possibility of multicollinearity affecting the estimates, we put more confidence in the reported results (that were obtained by using the largest possible number of observations.)[41]

6. Conclusion

In this paper a new, long and rich panel data set consisting of all Finnish publicly traded firms is assembled and then used to provide the most reliable evidence to date on a number of hypotheses relating to the adoption of selective versus broad-based stock options. As well as testing many standard hypotheses, that hitherto have only been examined using data for the U.S., we pay special attention to the hypothesis that firms with higher market value per employee will find it cheaper to provide equity incentives to a larger group of employees. Our analysis leads us to expect that stock option compensation should correlate positively with stock market developments. We also present a number of competing hypotheses, of which we pay especial attention to the hypothesis that the adoption of option compensation is expected to be related to changes in market value.

At the stock market level we find that general patterns concerning the adoption of option schemes correlate strongly with overall market developments. Stock option adoption is found to be a pro-cyclical phenomenon, becoming more common and inclusive during a stock market upturn, and less common and selective during a downturn. Specifically, option schemes were first introduced in Finland in the late 1980s, and after the deep depression they were revived again in 1994, after a particularly prosperous year in the stock market. The stock option boom coincided with the bull market of the late 1990s. In the years 1998-2000, broad-based stock options became very popular, especially in newly listed firms. However, after the stock market downturn the number of newly launched option schemes, especially broad-based ones, declined markedly.

These findings are consistent both with an explanation that stresses *changes* in market values as a primary explanatory variable as well as an explanation based on the importance of the *levels* of market values. However, firm-level analysis enables us to probe deeper concerning the two competing explanations. In this firm-level analysis, we find that higher market values per employee lead to higher probabilities of the adoption of both broad-based schemes and selective schemes, while higher returns predicts exclusively selective schemes.

We also test extant hypotheses concerning the impact of firm characteristics on the adoption of stock options. Often these findings corroborate those based on U.S. data. Thus consistent with previous results in the literature, our findings suggest that the adoption of broad-based plans is related to difficulties in monitoring employee performance. Another finding is that larger firms as well as firms with more dispersed ownership are more likely to adopt selective schemes. This finding is consistent with the hypothesis that option schemes are adopted in situations where the agency problem between owners and management is especially severe. Thus, our findings suggest that selective and broad-based schemes are aimed to solve different types of agency problems.

In other cases our findings differ from the U.S. literature. Thus we find that share returns from the past year affect the adoption of targeted stock options, but that there is no effect on broad-based plans. We find only weak evidence for a liquidity constraint hypothesis. Finally, we find no evidence for the hypothesis that foreign ownership influences the adoption of option schemes.

Since the hypotheses we develop and test are not specific to institutional settings (with the exception of the foreign-ownership variable which is found to be statistically insignificant), our findings for Finland may be expected to have more general applicability. Even the impact of differences in collective bargaining is potentially ambiguous. Typically labor unions do not negotiate about option schemes—so option adoption might be expected to be independent of collective bargaining. And when we turn to the European evidence (see Kalmi et al., 2004) stock options are found to have arisen in European countries with a variety of institutional set-ups—the Finnish experience with options is not exceptional when compared with these countries. In addition, while we have no direct evidence for other European countries for stock options, we note that in other related fields (such as effects of employee participation and profit sharing on business performance) the empirical findings that have emerged seem to apply in economies with very different institutional set ups.[42]

Since the use of option schemes is correlated with market values and stock market conditions, it is interesting to conjecture as to whether a similar increase in stock options will take place when the stock market revives. If not, and stock options prove to be a one-time management fad, will something else replace them during the next stock market upturn? Our sense is that the long-term importance of the stock options boom may have been that equity compensation instruments have been introduced in places where previously their broad use was rare, such as in Finland. Equity compensation, whether in form of options, restricted stock or other instruments, is likely to remain popular in listed firms.

Tables

Table 1: Change in HEX stock market indices

Year	General index	Portfolio index
1990	-0.380	n.a.
1991	-0.113	n.a.
1992	0.077	n.a.
1993	0.657	n.a.
1994	0.164	n.a.
1995	-0.062	n.a.
1996	0.411	0.322
1997	0.301	0.273
1998	0.524	0.138
1999	0.982	0.541
2000	-0.098	-0.242
2001	-0.367	-0.191
2002	-0.376	-0.150

Source: Helsinki Stock Exchange / Department of accounting and finance, Helsinki School of Economics

Notes: 1.The general index is weighted by firms' market values. Portfolio-index is calculated similarly, but the maximum weight assigned to one company is limited to 10%.

2. Entries represent changes from the previous year and are in logarithmic scale.

Table 2. The prevalence of stock option plans in Finland

Year	1.Nr of firms in the main list	2. Nr of firms in total	3. First option plan	4. First broad-based option plan	5. Nr of new plans in the main list (% of main list firms)	6. Nr of new broad-based plans in the main list (% of main list firms)	7.Nr of new option plans (all)	8. Nr of new broad-based option plans (all)	9. Nr of main list firms having option plans (% of main list firms)	10. Nr of main list firms having broad-based option plans (% of main list firms)	11. Nr of firms having option plan (% of all listed firms)	12. Nr of firms having broad-based option plan (% of all listed firms)
1987	52		1	0	1 (1.9%)	0	1	0	1 (1.9%)	0	1	0
1988	70		2	0	2 (2.9%)	0	2	0	3 (4.3%)	0	3	0
1989	82		4	1	5 (6.1%)	0	6	1	6 (7.3%)	0	7	1
1990	77		2	2	2 (2.6%)	1 (1.3%)	3	2	7 (9.1%)	1 (1.3%)	8	2
1991	66		3	0	4 (6.1%)	0	4	0	9 (13.6%)	1 (1.5%)	10	2
1992	65		1	0	1 (1.5%)	0	1	0	8 (12.3%)	1 (1.5%)	11	2
1993	60		4	0	5 (8.3%)	0	6	1	12 (20.0%)	1 (1.7%)	15	2
1994	68		20	2	16 (23.5%)	1 (1.5%)	21	2	27 (39.7%)	2 (2.9%)	34	3
1995	74		5	0	6 (8.1%)	1 (1.4%)	7	1	34 (45.9%)	2 (2.7%)	38	3
1996	73		3	2	7 (9.6%)	1 (1.4%)	9	3	34 (46.6%)	3 (4.1%)	36	6
1997	82	115	12	2	17 (20.7%)	3 (3.7%)	22	4	40 (48.8%)	4 (4.9%)	46 (40.0%)	7 (6.1%)
1998	92	119	24	14	37 (40.2%)	12 (13.0%)	47	17	60 (65.2%)	17 (18.5%)	69 (58.0%)	21 (17.6%)
1999	102	137	21	17	31 (30.4%)	15 (14.7%)	42	23	77 (75.5%)	30 (29.4%)	91 (66.4%)	36 (26.3%)
2000	107	150	20	16	44 (41.1%)	18 (16.8%)	61	30	88 (82.2%)	39 (36.4%)	113 (75.3%)	54 (36.0%)
2001	103	145	4	1	25 (24.3%)	7 (6.8%)	33	11	87 (84.5%)	38 (36.9%)	112 (77.2%)	54 (37.2%)
2002	99	137	1	0	22 (22.2%)	2 (2.0%)	33	6	82 (82.8%)	35 (35.3%)	101 (73.7%)	49 (35.8%)
2003	97	134	0	2	15 (15.5%)	2 (2.1%)	20	3	77 (79.4%)	33 (34.0%)	95 (70.9%)	44 (32.8%)
Total			127	59	240	63	318	104				

Source: Option database, Helsinki School of Economics. All our data on options presented in subsequent tables are from this source.

Table 3. Summary statistics

Variable	Mean	Standard deviation	Minimum	Maximum
Market value per employee (EUR)	206051.5	515397.6	4384.4	5629089
Share returns (ln)	0.006	0.559	-2.98	2.08
Annual volatility (std)	0.464	0.228	0.029	2.800
Foreign ownership (%)	17.87	20.36	0	97
Voting share of three largest owners (Ownership concentration) (%)	48.21	24.75	0.67	100
Net interest expenses / sales (Interest burden) (%)	1.17	2.97	-15.3	28.3
Number of employees	4920.57	8103.61	56	58708
Intangibles / fixed assets (%)	23.40	21.78	0.4	97.2
Previous option scheme	0.67	0.47	0	1

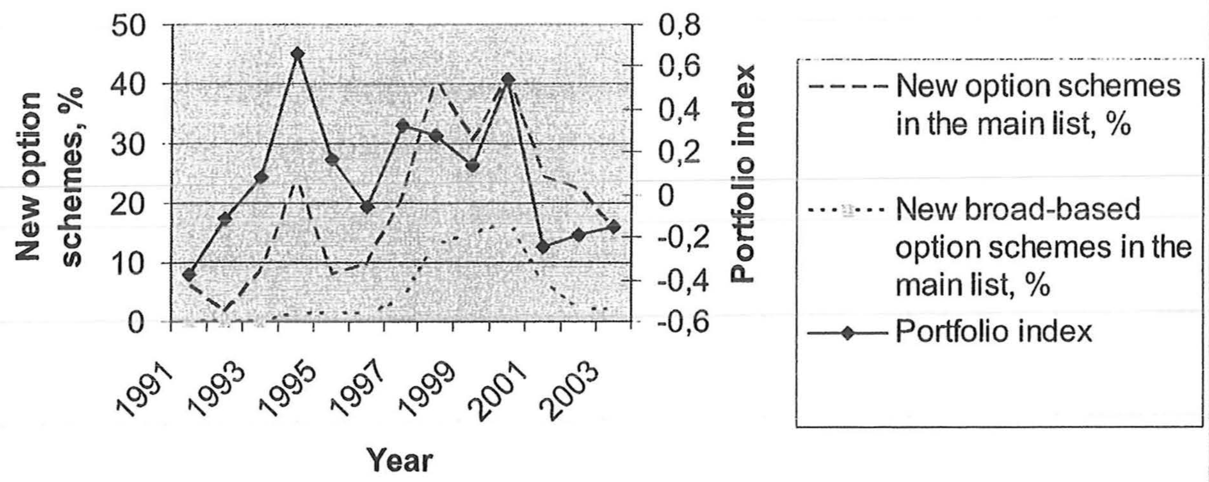
Note: 1. All monetary variables are measured in real terms.
2. Number of observations is 799.

Table 4. The determinants of option scheme adoption: multinomial logit models

	A1. Selective scheme	A2. Broad-based scheme	B1. Selective scheme	B2. Broad-based scheme	C1. Selective scheme	C2. Broad-based scheme	D1. Selective scheme	D2. Broad-based scheme
Market value per employee (ln)	0.049*** (3.20)	0.041*** (4.38)	0.044** (2.54)	0.026*** (3.05)	0.047*** (2.63)	0.021** (2.14)	0.043** (2.36)	0.015* (1.77)
Share returns (ln)	0.045*** (3.25)	0.004 (0.48)	0.050*** (3.61)	0.008 (1.11)	0.041*** (2.82)	0.000 (0.01)	0.036*** (2.60)	0.001 (0.24)
Annual volatility	-0.000 (-0.00)	0.016** (2.31)	-0.007 (-0.48)	0.006 (1.06)	-0.008 (-0.47)	0.009* (1.72)	0.001 (0.05)	0.004 (0.82)
Foreign ownership							0.005 (0.32)	0.001 (0.14)
Ownership concentration							-0.027* (-1.89)	-0.006 (-0.91)
Interest burden							0.024* (1.86)	-0.011 (-1.40)
Number of employees (ln)							0.057*** (3.34)	-0.001 (-0.08)
Intangibles / fixed assets							0.011 (0.68)	0.016** (2.40)
Previous option scheme							-0.012 (-0.38)	-0.021 (-1.32)
ICT & electronics			0.071 (1.12)	0.095 (1.44)	0.066 (1.09)	0.093 (1.52)	0.077 (1.20)	0.069 (1.18)
Manufacturing			0.040 (0.89)	0.004 (0.14)	0.039 (0.89)	0.005 (0.19)	0.019 (0.49)	0.004 (0.15)
Year dummies	NO	NO	NO	NO	YES	YES	YES	YES
Baseline probability	0.167	0.052	0.168	0.047	0.164	0.040	0.153	0.035
Wald Chi^2	83.15***		100.44***		146.96***		192.11***	
Pseudo R^2	0.075		0.091		0.111		0.152	

Notes: 1. Significance levels: * 10 %, ** 5 %, *** 1 %.
2. The reported coefficients denote the increase in probability of adoption when the explanatory variable increases by one standard deviation –unit (or changes from 0 to 1 for dummy variables). We report the z – statistics in parenthesis.
3. The number of observations is always 799.
4. The standard errors in multinomial logit model are corrected for clustering of firm-observations.

Figure 1. New option schemes in the main list and portfolio index



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Endnotes

[1] This paper has benefited from comments from participants at the 30th EARIE conference in Helsinki, August 2003, the ASSA conference in San Diego, January 2004, the IAFEP Conference in Halifax, July 2004, the EEA congress in Madrid, August 2004, and various seminars at the Helsinki School of Economics and Helsinki Center for Economic Research. The authors are extremely grateful to Seppo Ikäheimo for sharing his data on options and for his helpful comments, to Antti Kauhanen, Hannu Piekkola, Iikka Kuosa and Alexander Corporate Finance for sharing some of their data, to the Editor (David Levine), three anonymous referees, Uwe Jirjahn, Antti Kauhanen, Jeffrey Pliskin, Staffan Ringbom, Juuso Toikka and Otto Toivanen for helpful comments, and to Mikael Katajamäki for his outstanding research assistance. Kalmi and Mäkinen gratefully acknowledge a research grant from the LIIKE-program of Academy of Finland that has made this research possibly. In addition, Kalmi gratefully acknowledges financial support from the Marcus Wallenberg Foundation and the Helsinki School of Economics Research Foundation. Mäkinen thanks the Yrjö Jahnsson Foundation, the Helsinki School of Economics and the Foundation of Kluuvi for financial support. All support from the Research Institute of the Finnish Economy is gratefully acknowledged.

[2] In Finland, broad-based schemes always include managers and they are rarely egalitarian.

[3] The literature on performance effects has not reached definite conclusions on the effects of the stock options, especially on the merits of awarding broad-based stock options. Pessimists (notably Hall and Murphy, 2003) argue that the use of stock options imposes too much risk on ill-diversified employees. Furthermore, they argue that stock options do not improve incentives because employees are not able to effect the share price by their individual actions, and because employees are tempted to free-ride on the effort exercised by other employees. In contrasts, optimists (Sesil et al. 2002; Blasi et al. 2003) argue that stock options align the interest of the employees with those of shareholders, promote entrepreneurial behaviour among employees, help to retain the best employees, and are less risky than equity ownership.

[4] There have been two unpublished studies concerning the use of option schemes in Finland namely Pasternack (2002) and Pasternack and Rosenberg (2003). However these differ in important crucial respects from our approach. For example, Pasternack (2002) focuses on the adoption of the first option scheme and does not distinguish between managerial and broad-based schemes.

[5] There are other hypotheses on the reasons behind the popularity of options that we do not test in this paper, including the "accounting myopia" argument advanced by Hall and Murphy (2003) and the argument that stock options are the result of managerial power by Bebchuk and Fried (2003). Another, potentially very relevant motivation for stock option adoption are the issues related to recruitment and retention of employees. However, we cannot readily address these issues since we do not have relevant labour market data. In previous research, Oyer and Schaefer (2005) have used stock market data to proxy labour market developments. We cannot use the same approach since stock market developments are our main focus.

[6] The positive correlation between share returns and option use is posited also by Liang and Weisbenner (2001), who argue that options reward past performance.

[7] This hypothesis is supported by the observation by Holmström and Kaplan (2001, p. 140), who write that stock options were popular during the stock market boom of the 1960s, but they disappeared during the downturn in 1970s.

[8] The managerial power approach of Bebchuk and Fried (2003) would suggest an alternative hypothesis. It would predict that since managers are more weakly monitored under dispersed ownership, they are more likely to grant themselves options.

[9] For a more detailed exposition of corporate governance changes in Finland, see Hyytinen et al.(2003).

[10] At its peak, Nokia represented well over 50 % of the value of the stock exchange.

[11] Vartiainen (1998) provides a good overall presentation in English of the Finnish industrial relations system.

[12] According to a 1999 poll organised by Gallup Finland and commissioned by SAK (trade union confederation), only 4 % of the population accepts managerial option schemes uncritically and only 22 % think managers should be entitled to a substantially larger performance-pay component than other employees. 75% of respondents think that all employees should be entitled to stock options. Information taken from http://www.rakennusliitto.fi/press/gallup (26.3.2001)

[13] The resulting data are described in Ikäheimo et al. (2004). We are grateful to Professor Ikäheimo for giving us access to these data.

[14] We thank Antti Kauhanen and Hannu Piekkola for their help with these data.

[15] A description of these handbooks can be found in http://personal.inet.fi/yritys/porssitieto (30.11.2003). We thank Iikka Kuosa for pointing us out the data and providing some sample data he had collected.
[16] We do not possess data that would enable us to calculate the Black-Scholes values of options. However, we note that several studies highlight the poor applicability of these measures in the context of undiversified and risk-averse employees (e.g. Meulbrok, 2001; Hall and Murphy, 2002; Ikäheimo et al. 2005)
[17] Moreover, data on privately owned firms were not available.
[18] Our definition of broad-based schemes is thus different to some studies that have used a criterion that has been derived from firm-level surveys, such as 50% of the workforce (Kroumova and Sesil, forthcoming). An advantage of our measure is that it is derived from publicly reported sources that must be externally verifiable.
[19] Unfortunately, we cannot reliably distinguish in our data set whether the firm required this kind of investment from the employees or not.
[20] These data have been reported in Kalmi et al. (2004). We thank the project leader, Professor Erik Poutsma for allowing us to use these data in the comparisons.
[21] Notice that the respondents were asked to report their participation rates during spring 2001. The reported figure may differ from the participation rates in individual schemes, since the reported figure may be an outcome of several schemes. In this case, the figure may overestimate the participation rates in individual schemes. However, because of employee turnover (i.e., employees who subscribed options may no longer work for the company) the reported figure may also be an underestimate of participation in individual schemes.
[22] Among those classified as having broad-based schemes, the range of participation rates is from 30 % to 100 %, while among those classified as having selective schemes the range was from 1% to 25%.
[23] For instance, Beltton reports a 60% participation rate on its 2000 broad-based stock option programme, and Nokian Tyres report a 42 % participation rate on its 2001 programme. See www.beltton.com and www.nokiantyres.com
[24] Personal communication with Erkki Helaniemi, August 24, 2005. Mr. Helaniemi is a partner of Alexander Corporate Finance and has been personally in setting up dozens of Finnish stock option schemes.
[25] Dilution is defined here as the number of shares than can potentially be purchased by using the options, relative to the number of shares outstanding at the time of issue plus the number of shares available through options.
[26] However, the ratio of broad-based and selective schemes is likely to be substantially smaller in dilution than in participation rates, because lower-level employees are typically granted less shares than managerial employees.
[27] Other lists consist of firms that are of rather low economic significance compared to firms included on the main list. These smaller lists are maintained by investment banks and stock brokerage companies and are excluded from our analysis.
[28] Firms may adopt simultaneous schemes for many reasons. The firm may wants to broaden its schemes. Alternatively they may want to include new hires in a scheme. Or they may want a new scheme with conditions that better reflects the business prospects of the firm (for instance, if the scheme is far out of the money, the management may want to install a new scheme with lower exercise prices for incentive reasons).
[29] A firm is coded as having a scheme in year t if it has at least one scheme that has started in year t or earlier and if the final date for exercising options in this scheme is in year t+1 or later.
[30] We saw from Column 5 that new adoptions fell already after 2000, but since the option schemes are typically multi-year, the total number of option schemes reacts to the economic development with a lag.
[31] One interpretation of the data reported in Table 2 is that contagion may have played a role in the diffusion of options in Finland. In future work we plan to explore such issues. We believe that this is the appropriate course to follow (rather than address such matters in this paper) in part because matters surrounding contagion are difficult to pin down. For example, it is unclear whether contagion is predominantly by sector, by region, or by something else.
[32] This variable is lagged by one year since the decisions on option schemes are typically done in the shareholder annual meeting which is typically held during the spring, so the decision reflects the previous year's economic situation.

[33] This is, of course, not always possible since some variables are constructed to investigate hypotheses that have not yet been subject to empirical scrutiny (e.g. market value to employees and foreign ownership.)

[34] We exclude financial and real estate companies from the analysis. In the three first regressions presented in Table 4 the number of observations potentially was 853. However, 54 observations were dropped due to missing values in some variables in the full specification. To maintain comparability, we used the restricted sample in all estimations. The results would not change substantially, even if all the available observations were used.

[35] The multinomial logit method is sensitive to the problem known as "independence of irrelevant alternatives" (IIA), which means that the odds between any pair of alternatives do not depend on other outcomes that are available. This assumption may be problematic especially if two classes are perceived to be very similar to each other. We tested the IIA assumption formally by using the test suggested by Hausman and McFadden (1984). We implemented the test in STATA using the mlogtest command written by Long and Freeze (2003, p. 207-8). For the estimates reported here this test consistently indicated that the IIA assumption was not violated.

[36] The question of how best to proxy monitoring difficulties is a challenging one. In response to a suggestion from the editor, we experimented with approaches that use measures alternative to those reported in the paper including firm and industry Q (market value to book value of assets). These alternative measures were included in diverse specifications as were two other alternative measure, sales growth and wages, measures that have been used in previous studies to capture monitoring difficulties or human capital. While some specifications are characterized by strong multicollinearity between our proxies for monitoring and other key variables, we find consistent evidence that all our alternative proxies for monitoring difficulties have a statistically significant relationship with the adoption of broad-based stock option plans.

[37] This finding is also consistent with the share liquidity arguments of Holmström and Tirole (1993).

[38] This finding is consistent with a number of previous studies, including Core and Guay (2001), Ittner et al. (2003), and Kroumova and Sesil (forthcoming).

[39] Notice that the proxies for liquidity constraints have produced inconclusive results in prior research. While Core and Guay (2001) find support for the liquidity hypothesis, Ittner et al. (2003) and Kroumova and Sesil (forthcoming) do not.

[40] The results are available from authors upon request. Note also that we do not enter market value as such in any of the regressions. In HEX, there are other (small) firms where market value per employee has been even higher than in Nokia.

[41] These unreported results are available upon request from the authors.

[42] Compare for instance Sesil et al. (2003), Ichniowski et al. (1996), Handel and Levine (2004), Kato and Morishima (2002), Conyon and Freeman (2004), Eriksson (2003), and Bayo-Morines et al. (2003).

The Productivity Effects of Stock Option Schemes: Evidence from Finnish Panel Data[1]

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Abstract:
New panel data for all Finnish publicly listed firms during 1992-2002 enable us to distinguish broad-based stock option plans from those selectively allocated to particular employees. Whereas for broad-based schemes, we consistently find statistically insignificant associations with firm productivity, for selective schemes baseline estimators suggest positive and statistically significant productivity effects of between 2.1 and 2.4%. However, when endogeneity and dynamics are taken into account, no evidence is found of a link with firm productivity. Our main findings are consistent with hypotheses that predict negligible effects of option plans for enterprise performance, such as those based on free riding, or psychological expectancy theory or accounting myopia.

Keywords: Employee stock options; Compensation and compensation methods and their effects; Productivity; Panel data

JEL-codes: J33; M52; L25; C33

1. Introduction

During the 1990s, stock options became an increasingly popular compensation method in many countries (e.g. Hall, 1998; Murphy, 1999). Initially, option programs were typically allocated, almost without exception, only to executives.[2] But this association of stock options mainly with managerial compensation changed rapidly after more and more companies worldwide started to issue stock options to the workforce more broadly (e.g. Weeden et al., 1998; Lebow et al., 1998; Blasi et al., 2003). In turn this growth of stock options has generated heated public discussion with some viewing stock options as a device by which managers transfer excessive benefits to themselves, while others see options as a major innovation in managerial and personnel compensation.

The growth of options has also been accompanied by a mushrooming of theoretical and empirical literature on stock options. However, the economic impact of stock options remains a contentious issue, both theoretically and empirically. An important survey on the impact of executive equity compensation on firm performance found that "[T]here is presently no theoretical or empirical consensus how stock options and managerial equity ownership affect firm performance" (Core et al., 2003, p. 34). In contrast, existing empirical work on *broad-based* stock options consistently finds that firm performance is enhanced by stock options. This has led some observers to conclude that it is economically desirable to extend stock options to a large segment of employees (Blasi et al., 2003; Rosen, 2006). However, findings based on existing empirical work are potentially limited. For instance, empirical analysis of broad-based stock options have relied on survey data for a specific sector (which may not be representative), typically are for only short time periods and are mainly for the U.S. and the U.K. (e.g. Sesil et al., 2002; Ittner et al., 2003; Conyon and Freeman, 2004). Moreover, existing data may not enable a careful investigation of the productivity effects of different types of option plans.

By contrast in this paper we use new panel data that we have assembled. The data include all Finnish publicly listed firms during a relatively long period, namely 1992-2002. Thus it enables us to see if previous findings based mainly on evidence generated using US and UK data, are applicable to another country which once had a very different system of corporate governance but which has recently moved closer to an Anglo-Saxon model of corporate governance. In our empirical work we estimate Cobb-Douglas production functions with different stock option program indicators. Furthermore, whereas earlier empirical literature has used cross-section and fixed effects models, we also address the potentially important issue of endogeneity of inputs and options by estimating dynamic panel data models with a GMM estimator.

For broad-based option scheme indicators the key result is that different estimators consistently find a statistically insignificant association with firm productivity. For selective option schemes, the baseline fixed effects estimator suggests a 2.1-2.4% positive and statistically significant effect of the option plan size indicator on firm productivity. However, in empirical models in which endogeneity and dynamics are accounted for, no evidence is found of a link with firm productivity. Insofar as our findings do not provide support for hypotheses of a positive association between option schemes and firm productivity, our findings differ in important ways from earlier findings that are based on less rigorous methods and more limited data.

This paper is organised as follows. Section 2 provides the conceptual underpinnings for the study and also surveys relevant empirical research on the productivity effects of stock options and related schemes. In section 3 we describe the institutional framework and our most unusual data. Section 4 outlines our empirical strategy and this is followed by a presentation of our findings. In the final section we provide conclusions and implications of the paper.

2. Conceptual Framework and Previous Empirical Work

A key argument of proponents of stock options is that options can align the interest of employees and shareholders. For example, stock options may motivate employees to exert more effort and take actions that are mutually beneficial to both owners and employees. The motivational effect of stock options may be especially relevant in situations where alternative approaches, such as direct monitoring or piece rates, are not feasible, perhaps because of monitoring difficulties (Sesil et al., 2002).

Alternative arguments explaining the use of stock options are based on the idea that stock options increase employee total compensation in good times and decrease it in bad times. Stock options have been deemed crucial in recruiting and retaining employees, especially in markets where employees are potentially highly mobile (Rousseau and Shperling, 2003). Stock options may also help to retain key employees, both because they adjust pay according to current market conditions (Oyer, 2004) and because they are a deferred form of compensation.

Other papers view stock options as a substitute for fixed wage contracts. Inderst and Müller (2005) show that stock options may prevent inefficient firm closures, since they substitute for fixed wage contractual payments and hence save firms cash expenses in bad times. Arya and Mittendorf (2005) argue that when a manager is willing to accept stock options as a part of her compensation package (and, by assumption, a lower fixed wage), she is indicating her confidence in her ability to raise firm value. Thus stock options provide a useful signalling device of managerial quality.

However, there has also been criticism of stock options. For example, stock options, when exercised, entail a cost to shareholders in the form of dilution of ownership. Also others note that the costs of stock options have not been included in income statements (e.g. Hall and Murphy, 2003). Hence they argue that the increasing popularity of options in part reflects firms' mistakenly thinking that options are a cheaper form of compensation than their true cost.

Also, a potential performance impact has been questioned. Payment schemes that reward collective performance suffer from the free-rider problem: an individual who increases his effort will bear the full cost of the increase in effort, but will realise only a small part of the resulting increase in output (e.g. Alchian and Demsetz, 1972; Oyer, 2004). Another criticism comes from psychological expectancy theory (Vroom, 1995). According to the "line-of-sight" argument, rewards based on performance can only be motivating if, by their actions, employees can influence the measures on which performance-pay is based. This is typically not the case with stock option plans, where employees (with possible exception of top executives) can hardly perceive any direct link between their actions and the share price performance.

Both free-rider and line-of-sight arguments have been countered in the literature. First, when rewards are based on group performance, according to Kandel and Lazear (1992) it is in the interest of individual employees to develop a group norm where the employees monitor the performance of their peers and prevent free-riding behaviour. Second, equity schemes may create a common bond or "psychological ownership" among employees and thus change their behaviour so that it would match the collective interest (Pierce et al., 1991; Baron and Kreps, 1999).

Another cost related to equity schemes is that they increase the risk employees are facing, since employees have both a substantial proportion of their financial capital and human capital invested in one workplace. Since they do not value their options to the same extent as an outsider would, employees may require higher total compensation (Meulbrok, 2001). However, there can be important differences between executives and employees in these respects. In the compensation of top executives, the line-of-sight and diversification problems are not as severe as with lower-level employees (Hall and Murphy, 2003). On the other hand, executive stock option compensation may be motivated by rent-seeking activities (Bebchuk and Fried, 2003).

Ultimately the impact of stock options on business performance is an empirical question. Most work has concentrated on the performance consequences of executive stock options. For instance, by using Execucomp data, Hanlon et al. (2003) find that changes in Black-Scholes values of option grants are positively associated with future operating income of the firm. However, using similar data, Larcker (2003) finds that results depend crucially on estimation strategy. Core et al. (2003) conclude that the research consensus is that there is no clear connection between executive equity compensation and firm performance.

The empirical findings concerning broad-based stock options have tended to be markedly more positive, though the literature is substantially smaller. Conyon and Freeman (2004) examine the economic outcomes of broad-based option schemes in a sample of UK listed firms during 1995-1998. They use three survey data sets for 284 firms, estimate fixed effects regressions and find evidence that the presence of a stock option plan is significantly associated with higher firm-level productivity. Sesil et al. (2000) use survey data on broad-based option schemes from different sectors in the US provided by the National Center for Employee Ownership. They find that, in 1997, stock option firms had 28% higher productivity than non-option firms and 31% higher productivity than their non-option pairs. However, the response rate of the survey is only 10% yielding 73 firms in the final data set. In a subsequent study the same four authors (Sesil et al., 2002) compare the performance of 229 new economy firms, which offer broad-based option schemes to their non-stock option granting counterparts. They find evidence that productivity is higher in firms with broad-based plans. Their research methods include descriptive analysis, paired matching comparisons between broad-based and non-broad-based options firms within the same industry, and a cross-section regression for 1997. Ittner et al. (2003) use survey data to examine 217 new economy U.S. firms during 1999-2000. By using cross-section analyses they examine the performance consequences of option and equity grants to senior-level executives, lower-level managers, and other employees. Their findings indicate that lower than expected option grants and/or existing option holdings are associated with lower accounting and stock price performance in subsequent years.

In sum, the empirical evidence from these studies suggests the existence of a positive and often quite sizeable link between broad-based stock option plans and productivity at the firm-level. In turn, this implies that the available evidence provides support for theorists who predict that potentially powerful economic effects will flow from options and dominate the effect of factors such as free riding, accounting myopia and managerial rent-seeking. However, before accepting this conclusion it is important to note some key shortcomings of these studies. For one thing, in all of these studies, often the data are based on surveys and are apt to suffer from various selection biases, for example three studies focus only on new economy firms. Second, nearly all studies are limited insofar as they concentrate on the time-period before the stock market collapse in 2000. Third, and at odds with most theory, no studies are able to reliably distinguish the productivity impact of selective versus broad-based plans. Fourth, the econometric methods that the available data enable researchers to use are sometimes less than desirable. Thus some studies have access only to cross-sectional data, and few appear to seriously address the potentially highly important issues of endogeneity of option schemes and inputs used in a production process.

While there are only a small number of empirical studies on the impact of stock options on productivity, the empirical literature on the productivity effects of other forms of employee financial compensation that are alternative to the traditional fixed-wage arrangements, such as *employee profit-sharing* and *employee stock ownership plans (ESOPs)* is quite large.[3] Consequently, it is useful to briefly highlight some key issues and findings in that literature since this may help to shape the empirical strategy we adopt in this study.

Typically studies of firms with *employee profit-sharing plans* find a positive relationship between profit-sharing and firm productivity. This is the conclusion of several surveys including, for example, Weitzman and Kruse (1990) and Jones and Pliskin (1991). This typical finding emerges from the empirical studies that employ diverse methods to investigate profit sharing arrangements that exist in a variety of institutional settings including the former West Germany (Cable and Wilson, 1990), the UK (Wadhwani and Wall, 1990), the US (Kruse, 1992) and Finland (Kauhanen and Piekkola, 2002). Findings based on studies of firms with *employee stock ownership* *plans* also typically support the existence of a positive relationship between ESOPs and firm productivity or performance. However, as many surveys point out (e.g. Kruse, 2002), the evidence in support of this positive link is probably less robust than for profit sharing. Again there is evidence that employee stock ownership can be positively associated with enhanced business performance in a variety of institutional settings including Japan (Jones and Kato, 1995) and the U.S. (Kumbhakar and Dunbar, 1993).

However, this other literature also draws attention to the potential sensitivity of findings to several factors. Prominent among these is the need to have data that are not distinguished by various kinds of selection bias, a common problem with survey data. Another key matter is the issue of institutional detail—the form of the ESOP or profit sharing arrangement often matters. Also, some theorists argue that for sustained effects on enterprise performance, financial participation must be accompanied by changes in decision-making participation. Hence a failure to include controls for such other factors may lead to empirical models that are misspecified (e.g. Conte and Svejnar, 1988). Finally, there is the issue of the appropriate econometric approach. Amongst several potential matters, the sensitivity of findings to potential issues of endogeneity of plan schemes and inputs used in production is clear.

In devising our empirical strategy we will respond to these issues that are highlighted by the concerns as best we can. Our data will enable us to address most of these matters. The only exception is our lacking data for other HR practices such as participation in decision-making. Because we utilise public firm-level data on stock option plans, we cannot control for the level of employee participation in decision-making, the existence of profit sharing, and other human resource management practices. We can, however, separate firm fixed effects and common time-specific effects from several other factors that possibly have effects on firm productivity by using fixed effects and GMM estimators.

3. Institutions and the Data

In this section we describe the institutional context and the data. We examine the impact of stock option compensation on firm productivity by assembling new panel data for Finnish firms for 1992-2002. This was a particularly turbulent period in Finnish economic history. In 1990, Finland had just entered a deep depression, which was the most severe of any OECD country since the Second World War. In 1995, Finland joined the European Union and, in 2002, adopted the common European currency, the Euro, in the first wave of adoptions.

In industrial relations, the most marked change during the period of interest was the increased use of performance-related pay (Kauhanen and Piekkola, 2002). On the other hand, collective bargaining and centralised income agreements remained intact. The unionisation rate has remained relatively high, between 70 and 80 per cent.

The increase in stock option compensation reflects a deep change in the Finnish corporate governance system. In the end of 1980s, the Finnish corporate governance system in listed firms was very much bank-centred and resembled the German system. The stock market started its recovery after the recession in 1993, and the importance of the equity market in financial intermediation grew throughout the 1990s. Both the turnover and market value of firms listed on the stock exchange increased dramatically throughout the decade, with Nokia leading this development. In the end of the1990s the Helsinki Stock Exchange saw a wave of new listings.

Now stock markets are much thicker, more transparent and arguably provide more reliable information than in the past. At the same time, both monitoring of insider trading and legal punishments have become stricter. During the last 10-15 years Finland has shifted from a bank-based financial intermediation closer to a market-based Anglo-American system. As part of this institutional change publicly listed Finnish firms have adopted stock option schemes extensively in the 1990s. As discussed below, the most active period of stock option adoption coincided with the height of the stock market boom in the late 1990s. However, as stock market prices started to fall after May 2000, accelerating further in 2001 and 2002, the rate of stock option adoption decreased markedly.

All of our firms are traded on the Helsinki Stock Exchange (HEX). However, firms that were on the two smaller lists (i.e. the Over-the-counter and the Stockbroker's list), that were maintained by investment banks and stock brokerage companies, are excluded before 1997 due to their rather low economic significance compared to the main list. Since 1997 HEX has taken over the smaller lists and also has started to operate two additional lists besides its main list: the "I" (Investor) -list and the "NM" (New Market)-list. The "I-list" consists of firms that are traded infrequently and are often majority-owned by large investors. The "NM" list consists of smaller IT and high technology firms, similar to the NASDAQ or the Neuer Markt in Frankfurt. Thus, we have information on the presence of option schemes on the main list throughout the period and on the minor lists since 1997. However, we do not have option program information on firms that have not been listed in the HEX, since our option data are based on public information of listed firms. We are aware that some unlisted Finnish firms have adopted option schemes, at least during the bull market at the end of the

1990s. Unfortunately, there was no option data information available for these firms.[4] We expect that these programs were more likely to be located within the ICT sector than in other sectors. We believe, however, that the set of these unlisted firms is moderate, since stock option compensation works properly only in situations where the value of shares can be assessed by the stock market.

[Table 1 about here]

Our panel data on stock options were initially organised by Professor Seppo Ikäheimo from the Helsinki School of Economics. However we have used several sources to complement and update the original data. These include annual reports, stock market releases and option data obtained from Alexander Corporate Finance, an investment bank. We use these data to briefly describe the general evolution of Finnish stock option plans during 1992-2002.[5]

Column (1) in Table 1 gives the total number of firms during the period. The number of firms at the HEX fluctuates considerably, which partially relates to the business cycle. Column (2) describes the number of new option plans. The early peak year is 1994 (21 new plans). Then the number increases from 1997 (22) until 2000 (61). Thereafter it drops to 33 for 2001 and 2002. Column (3) shows the development of new broad-based plans. The introduction of such schemes is concentrated during the years 1999-2000, when one-half of new plans were broad-based. In Columns (4) and (5) we approach this issue from another angle and provide time series data on the existence of option schemes. In these columns, we also use information on the timing of the scheme as well as on the launching of the scheme. In Column (4), a quarter of listed firms had an option scheme in 1993. This proportion jumps to around 50% in 1994, where it stays until 1996. After a temporary drop in 1997, the proportion increases from 1998 (58%) until 2001 (77%). In 2002 74% of listed firms had an existing scheme. Column (5) shows the development for broad-based plans. Roughly 3-8% of listed firms had an existing broad-based scheme in 1992-1997. This fraction steadily increases until 1998-1999, and stays around 36% in 2000-2002.

By combining the option data set with firm-level financial statements obtained from Balance Consulting, a consulting firm, we assemble firm-level panel data for 117 publicly listed firms from 1992 to 2002.[6] For each firm there are between four and eleven observations. Finally, we deflate all our nominal monetary variables to real euros for 2000 by using industry-specific gross output deflators, published by Statistics Finland. Table 2 summarises the pattern of our panel data.

[Table 2 about here]

In the analysis that follows, we distinguish between broad-based and selective schemes. The latter are mostly managerial schemes, although they can also include other key personnel (e.g. R&D workers). However, in order to qualify as a broad-based scheme, all employees (or at least a great majority) should be eligible. The classification is based on public stock exchange reports.[7] Finnish Law on Joint Stock Companies requires firms to report all relevant conditions about stock option schemes to shareholders prior to adoption. While a high rate of eligibility does not automatically guarantee a high participation rate, there are good reasons to believe that these are closely connected. For one thing, employees usually face only small costs when they subscribe to options – e.g. by providing a zero-interest loan to the company, with the company repaying the loan at face value after a certain period, usually 1-3 years. Thus, while employees face a cost in terms of foregone interest and liquidity, typically this cost is far below the real value of the options. Moreover, not all companies use this procedure, but rather they essentially give options to employees for free.[8]

In our empirical work, we develop three option program indicators. These measures reflect the presence or absence of an option scheme, the size of the scheme and whether the scheme is selective or broad-based. Two measures are binary variables and one is a continuous variable.[9] Our *first binary indicator* is *opt* measuring the presence of a scheme in a firm in given year t. It equals one for the group of option firms and zero otherwise. Thus, the indicator distinguishes option and non-option firms allowing us to examine the average impact of the presence of options on firm productivity.

Our *second binary indicator* also measures the presence or absence of a plan but it distinguishes between selective *(ssopt)* and broad-based *(bbsopt)* option plans. By a selective plan we mean a scheme that is targeted to a selected group of employees. These schemes include managerial programs, but also schemes that are targeted to key personnel. Broad-based plans are all encompassing, including managers, but they do not have to be egalitarian in the sense of all participants having the same number of options. By using these distinct dummy variables we can examine whether the average impact of plans on firm productivity differs between selective and broad-based option schemes.

Our *third program indicator* is potential dilution *(dilu)*. This indicator measures the potential size of effective schemes in firm *i* in year *t*.[10] This is a continuous variable – the ratio of the number of shares that may be awarded through effective stock option plans in a given year divided by the sum of total number of shares and the number of new shares that may be awarded through options at the end of a year. If a program ends in the middle of the year *t*, then the year *t-1* is the last year used in calculating dilution. The indicator distinguishes option and non-option firms allowing us to examine the average impact of options on firm productivity. To investigate whether the productivity impact varies by plan characteristics, we also use separate dilution indicators, namely *diluss* for selective and *dilubb* for broad-based plans. To capture possible dynamic effects of a program, we also use once lagged dilution indicators.

It is worth stressing that our panel data include almost all listed Finnish companies during the period 1992-2002. We exclude only a few firms with less than four consecutive observations. This is mainly because there is some entry and attrition of listed firms at the Helsinki Stock Exchange (HEX). Also, some firms merged during the period. In this case, we have included only merged firms and excluded all information prior to the merger. Also, we exclude a firm if data for a key variable such as value-added are missing. Finally, to exclude potential outliers, we delete observations where: employment is less than 50 (32 firm-year observations); fixed capital is less than €1,000,000 (23 firm-year observations); or employment is more than 50,000 (4 firm-year observations). Table 3 presents summary statistics.

[Table 3 about here]

Table 4 presents the key variables grouped by a firm's option program adoption status. We observe that option firms have higher value added, bigger labour forces and they also use more fixed capital compared to firms without stock option schemes. For example, the mean value added for selective scheme firms is 496 million euros, whereas for broad-based firms it is 166 million euros and for non-option firms only 105 million euros. Table 4 also shows that large Finnish firms have preferred targeted schemes to broad-based option schemes. Finally, the mean value added per employee is about 3.4% higher in selective scheme firms than in broad-based firms (57,064 euros compared to 55,205 euros).

[Table 4 about here]

4. Econometric Strategy

We test two key hypotheses, namely: (i) that firm-level productivity is expected to be higher in option than in non-option firms; (ii) that the impact of options on firm productivity is expected to be dependent upon whether the plan is broad-based or selective. Our basic empirical strategy is to use a production function approach and panel data estimators. First, we estimate a series of baseline fixed effects estimators by assuming that all explanatory variables are strictly exogenous. Second, we estimate dynamic panel data GMM estimators to account for the potential endogeneity of a firm's decisions on inputs and option schemes. The following issues have influenced the specific empirical strategy we adopt.

First, we assume a Cobb-Douglas form of technology, since it has been used frequently in the related literature such as the evaluation of the effects of ESOPs on firm productivity (e.g. Jones and Kato, 1995) and when analysing the effects of stock options on firm performance (e.g. Conyon and Freeman, 2004.) Second, although the Cobb-Douglas functional form is more restrictive than other functional forms such as the translog, we prefer the Cobb-Douglas production function since, when accounting for endogeneity of inputs and options in GMM models, the instrument matrix Z_i may become sizeable under the translog specification, thereby biasing estimates in finite samples.[11] Third, we assume that option schemes may only have a direct impact on firm productivity. Fourth, since we do not have information on the detailed terms of option schemes, such as the exercise prices of options, we must bypass potentially important matters surrounding this issue.[12]

There are two reasons for using the fixed effects estimator in our baseline estimates. In part this is pragmatic – we use the fixed effects estimator because it has been used in the previous studies. For example, the estimator has been used in assessing the productivity effects of ESOPs (e.g. Jones and Kato, 1995) and stock options (e.g. Conyon and Freeman, 2004). Second, as is well known, firm fixed effects allow us to control for unobserved time-invariant differences in firms, such as managerial ability, employee quality and organization structure. We denote a firm's production function by $f(.)$, which relates firm value added[13] at time t, i.e. va_{it}, to inputs used in production and control variables:

(1) $va_{it} = f(k_{it}, l_{it}, eo_{it}, x_{it}, \eta_i; \tilde{\beta})$, where i=1,2, ..., N and t=1,2,...,T.

In Equation (1) subscripts i and t index firm and time, respectively. Firm deflated fixed capital is k_{it}, the sum of a firm's tangible and intangible assets at the end of the year, and labour input l_{it} is the mean number of employees in a given year. The option program indicator is denoted by eo_{it}, and x_{it} is a vector of control variables including industry-specific year dummies (for the ICT, the manufacturing and the service sectors) to control for industry-specific technological changes and economic shocks.[14] By η_i we control for unobserved heterogeneity among firms. The vector of parameters is $\tilde{\beta}$, and we are interested in the parameters for capital, labour and the option program indicator, i.e. $\beta_k, \beta_l, \beta_{eo}$. A baseline fixed effects specification for Equation (1) is the following Cobb-Douglas production function:

$$\text{(2)} \quad \begin{aligned} &\ln va_{it} = \beta_k \ln k_{it} + \beta_l \ln l_{it} + \beta_{eo} eo_{it} + \beta_x x_{it} + \eta_i + \varepsilon_{it}, \text{ where} \\ &\varepsilon_{it} \sim iid(0, \sigma^2);\ \text{i=1,2,..., N; t=1,...,T.} \end{aligned}$$

In Equation (2) the variables are the same as in Equation (1). Thus, a firm's inputs in the production process are capital k_{it} and labor l_{it}; eo_{it} is an option program indicator; x_{it} is a vector of possible control variables including industry-specific year dummies; η_i's are individual firm fixed effects and ε_{it} is the error term. The estimation results for Equation (2) are reported in section 5.

If the assumption of strict exogeneity assumption is violated, the baseline fixed effects estimator is potentially inconsistent. Therefore, we relax the strict exogeneity assumption on capital and labor inputs as well as on option schemes and estimate dynamic GMM models.[15] We also use these models to explore the dynamic effects of stock options. To obtain asymptotically consistent parameter estimates when an explanatory variable is likely to have violated the strict exogeneity assumption, we estimate single equation GMM estimators[16] by assuming the following dynamic model:

$$\text{(3)} \quad \begin{aligned} &\ln va_{it} = \beta_{va} \ln va_{i,t-1} + \beta_k \ln k_{it} + \beta_{k\_1} \ln k_{i,t-1} + \beta_l \ln l_{it} + \beta_{l-1} \ln l_{i,t-1} + \beta_{eo} eo_{it} \\ &+ \beta_x x_{it} + \varepsilon_{it}, \text{ where } \varepsilon_{it} = \eta_i + v_{it};\ v_{it} \sim \text{iid}(0,\sigma^2);\ \text{i=1,2,...,N; t=2,3,...,T.} \end{aligned}$$

In Equation (3) all variables correspond to those in equation (2). The presence of individual effects η_i in the error term ε_{it} implies that the lagged dependent variable $va_{i,t-1}$ is positively correlated with ε_{it}. Thus, at least in large samples with serially uncorrelated error terms v_{it}, it can be shown that the OLS level estimator for β_{va} is inconsistent. Furthermore, the omitted variable literature implies that the OLS level estimator for β_{va} is biased upward in large samples (e.g., Bond, 2002).

The fixed effects estimator (the within group estimator) removes this inconsistency by transforming each variable to be its deviation from its firm mean. However, if the number of time periods is small, the within group transformation introduces a non-negligible negative correlation between a transformed lagged value added $va_{i,t-1}$ and a transformed error term v_{it}. This result indicates that, at least in large samples (N large), the fixed effects estimator for β_{va} is biased downward (e.g. Bond, 2002).

The fact that the OLS level estimator for Equation (3) is likely to be biased upwards and the fixed effects estimator is likely to be biased downwards can be useful information in assessing whether an estimator is consistent. In other words, a consistent estimator would lie between the OLS level and the fixed effects estimator. If we do not observe this pattern or an estimator is close to either the OLS level or the fixed effects estimator, we might suspect severe finite sample bias or inconsistency (e.g. Bond, 2002).

We follow a dynamic panel data GMM estimation strategy and allow explanatory variables to be correlated with the individual effects η_i, since we exclude these effects from Equation (3) by a first-difference transformation:

$$\begin{aligned} &\Delta \ln va_{it} = \beta_{va}\Delta \ln va_{i,t-1} + \beta_k \Delta \ln k_{it} + \beta_{k\_1}\Delta \ln k_{i,t-1} + \beta_l \Delta \ln l_{it} \\ (4)\quad &+\beta_{l\_1}\Delta \ln l_{i,t-1} + \beta_{eo}\Delta eo_{it} + \beta_x \Delta x_{it} + \Delta v_{it}, \\ &\text{where } |\beta_{va}| < 1;\ i=1,2,...N;\ t=3,4,...T. \end{aligned}$$

In Equation (4) we assume that the initial conditions va_{i1} are predetermined, i.e. they are uncorrelated with the subsequent error terms v_{it}, $t = 2,3,...,T$ and that the error term v_{it} is serially uncorrelated. We then apply the lagged levels of $va_{i,t-1}$ dated at t-2 and t-3 as instruments for the corresponding first-differenced variables.

To account for potential endogeneity of inputs and options in equation (4), we proceed in two steps. In the first stage, we assume that only labour and capital inputs are endogenous variables. In the second stage, we also assume that option schemes are endogenous (as well as the input variables.) When addressing the potential endogeneity of capital k_{it} and labour l_{it} inputs, we assume that k_{it}, l_{it} are predetermined and use the lagged levels of k_{it} and l_{it} dated at t-1 and t-2 as instruments for the corresponding first-differenced variables. In other words, we assume that there is no contemporaneous correlation between the inputs and the error term v_{it}, but that the inputs may be correlated with $v_{i,t-1}$ and earlier shocks.[17] To address endogeneity of option schemes, we treat the dilution option program indicator $eo_{i,t}$ as predetermined.[18] Then we use the once lagged variable, denoted t-1, as an instrument for the corresponding first-differenced option program variables.

By accepting the moment condition restrictions above, we may construct an instrument variable matrix Z_i, where the lagged levels of explanatory variables are used as instruments for the corresponding first-differenced variables.[19] We follow the terminology suggested by Bond (2002) and call these estimators the differenced GMM estimators. As an extended estimator we apply the system GMM estimator[20] by also assuming that the levels of explanatory variables, i.e. k_{it}, l_{it} and $eo_{i,t}$, are uncorrelated with individual effects η_i and predetermined with respect to the error term v_{it}.[21] Thus we use lagged first-differences of $k_{i,t}, l_{i,t}$ and $eo_{i,t}$ as instruments for the GMM level equations. The estimation results for the differenced and the system GMM estimators are separately reported in section 5, under both assumptions concerning endogeneity.

5. Empirical Results

Table 5 reports the baseline contemporaneous fixed effects estimates for a Cobb-Douglas production function during 1992-2002.[22] The estimator deviates from the standard fixed effects estimator in that we have specified a first-order autocorrelation process in the residuals. This is the preferred estimation approach, since the autocorrelation tests strongly indicates that the disturbance term is first-order autoregressive.[23]

Three main conclusions emerge from Table 5. First, the estimates for capital and labor are highly significant in columns (1)–(4). The baseline elasticity of capital input is close to 0.15, whereas for labor it is about 0.62.[24] Second, in columns (1) and (2), where our option program indicator is the presence of a plan, we do not find statistically significant evidence of contemporaneous association between options and firm productivity. In column (1) the parameter estimate for the option program indicator is 0.002, but it is statistically insignificant. In column (2) the signs of parameters differ between selective and broad-based indicators. The selective scheme estimate is 0.015 and the broad-based scheme -0.028. However, both are statistically insignificant. Third, in columns (3) and (4), where our option program indicator is the size of a plan, we find statistically significant evidence of contemporaneous association between selective schemes and firm productivity (at the 10% level.) The parameter estimate for selective schemes is 0.84, whereas for broad-based schemes it is statistically insignificant -0.256. The mean dilution for selective schemes is 0.0286 indicating, on average, a 2.4% effect on firm productivity (0.0286*0.84=0.024).

[Table 5 about here]

Tables 6-8 show estimation results for the OLS level, the fixed effects, the differenced GMM and the system GMM estimators for a Cobb-Douglas production function.[25] The reported GMM estimates are based on the two-step GMM estimator with heteroskedastic-consistent asymptotic standard errors.[26] We also perform a finite-sample correction proposed by Windmeijer (2000), since simulation studies have shown that these standard errors are downward biased.[27] For all test statistics, we rely on the two-step GMM estimator.

In Table 6 we relax the strict exogeneity assumptions on inputs. Our program indicators are dummy variables, i.e. we measure the presence of a plan. The following key findings emerge from Table 6. First, the OLS level parameter estimates for va_{t-1} in columns (1) and (2) are substantially higher than the fixed effects estimates in columns (3) and (4). As noted earlier, a consistent GMM estimator would lie between these two estimators. Unfortunately we find that the differenced GMM estimates for va_{t-1} in columns (5) and (6) are below the fixed effects estimates. Thus we suspect severe finite sample bias or inconsistency which, in this case, is likely to be associated with weak instruments for individual series that are highly time persistent.[28] Another indication of inconsistency is that the differenced GMM parameter estimates for capital inputs reported in columns (5) and (6) are about twice as large as the OLS level and the fixed effects estimates reported in columns (1)-(4).

Second, Table 6 suggests that the estimates for the presence of a plan are statistically insignificant. In our preferred specifications, reported in columns (7) and (8), where we have controlled for simultaneity of inputs by treating them as predetermined, the signs of the option and selective scheme indicators are positive, but the coefficient for the broad-based scheme indicator is negative. In sum, we do not find statistical evidence of a contemporaneous association between option programs and firm productivity.

Third, the system GMM parameter estimates using a lagged dependent variable and reported in columns (7) and (8) are lower than the OLS level estimates but higher than the fixed effects estimates. This finding indicates that the system GMM estimator is likely to be consistent, at least for the lagged dependent variable.

Fourth, the autocorrelation tests, namely m1 and m2 reported in columns (7) and (8), provide support for the system GMM estimator. The tests indicate significant negative autocorrelation in the first-differenced residuals but not in the second-order residuals. This is exactly how it should be, if the disturbances are serially uncorrelated, indicating that the key assumption for the consistency of the system GMM estimator is fulfilled. Moreover, the Sargan tests clearly fail to reject the over-identification restrictions in columns (5)-(8), supporting the hypothesis of the validity of the instruments.

[Table 6 about here]

Next we focus on the endogeneity of an option program, since that may be driving the baseline fixed effects estimates reported in Table 6. The findings reported in Table 7 are based on a program's size indicator[29], but otherwise the estimation approach is similar to that underlying the findings presented in Table 6. Since the OLS level and the fixed effects estimates for the lagged dependent variable, the capital and labor inputs are almost the same in both the tables, we do not discuss these findings any further.

The following key findings emerge from Table 7. First, the fixed effect estimate for selective programs reported in column (4) support the positive association, reported previously in Table 6. Now the parameter estimate is 0.73 and it is statistically significant at the 10% level. The mean dilution for selective schemes is 0.0286 indicating, on average, a 2.1% effect on firm productivity (0.0286*0.73=0.021). Note, however, that the fixed effects findings in columns (3) and (4) are based on the assumption that the explanatory variables are strictly exogenous.

Second, the system GMM estimators in columns (7)-(10) suggest that, after controlling for potential endogeneity of the explanatory variables, all the estimated option dilution indicators are found to be statistically insignificant. In columns (7) and (8), where we have controlled for simultaneity of capital and labor inputs by treating them as predetermined, the signs of all indicators are positive. In columns (9) and (10) we also treat the dilution indicators as predetermined (to control for simultaneity), but even then we do not find any evidence that is statistically significant that programs can be associated with firm productivity. The parameter estimate for the selective scheme reported in column (10) is now about one third as large (0.25) as the statistically significant fixed effects estimate of 0.73 reported in column (4). The parameter estimates for the broad-based dilution indicator (column 10) is -0.542. In sum, after controlling for endogeneity, at conventional levels of statistical significance, we do not find any evidence that option programs affect firm productivity. This conclusion holds even in estimates that distinguish selective and broad-based option schemes.[30]

[Table 7 about here]

In Table 8 we expand our investigation to account for the dynamic effects of option programs, since a stock option program typically spans several years. Hence the effect on productivity may be realized with a lag. To account for dynamics we keep all the assumptions used in the models reported in Table 7, but in Table 8 we report estimates that use both contemporaneous and once lagged program indicators. As an option program indicator we use the size of a plan. A Wald test is used to determine whether contemporaneous and lagged indicators are jointly zero. We first estimate the OLS level and the fixed effects models in columns (1)-(4), thereafter the system GMM estimators in columns (5)-(8).[31] In columns (5) and (6) we treat capital and labor inputs as predetermined, and in columns (7) and (8), besides capital and labor, we also treat the program indicator as predetermined.

The key finding is that we do not find convincing statistical evidence of an association between option programs and firm productivity. The evidence for a selective scheme having positive effects on productivity when using fixed effects estimators and reported in Tables 5 and 7, is not supported in the dynamic models. While the sum of contemporaneous and lagged parameter estimates of selective schemes is 0.81, they are jointly statistically insignificant (p-value 0.24). Also, the system GMM estimates for program indicators are all found to be statistically insignificant.

6. Conclusions and Implications

In this paper we assemble new panel data for all Finnish publicly listed firms during a relatively long period, namely 1992-2002. Our data enable us to distinguish different types of option plans and to seriously address issues of endogeneity concerning options and inputs. Consequently we are able to see if previous findings that are based mainly on evidence generated using data that are less representative and for shorter time periods are sustained.

We proceed by estimating Cobb-Douglas production functions with three different option program indicators. These measures reflect the presence or absence of an option scheme, the size of the scheme and whether the scheme is selective or broad based. Furthermore, the long panel nature of our data allows us to estimate dynamic panel data models with a GMM estimator and thus address the potentially important issue of endogeneity of inputs and options.

The most important finding, yielded almost consistently in diverse specifications, is a statistically insignificant association between option programs and firm productivity. This result is exceptionally robust for broad-based schemes and is independent of what option program indicator is used in estimations. As such our findings are consistent with those who hypothesize that the performance impact of options will be limited because of reasons such as free-rider problems (e.g. Oyer, 2004), accounting myopia (e.g. Hall and Murphy, 2003) or line-of-sight arguments (e.g. Vroom, 1995). As such our results are consistent with much of the financial literature that does not find evidence of a link between options and business performance (e.g. Hall and Murphy, 2003.)

For selective programs, however, findings are less consistent. In our baseline fixed effects estimates we find a statistically significant productivity impact that is between 2.1-2.4%. Since most selective plans are allocated to executives and/or key employees, this finding also provides support for the line-of-sight argument – rewards based on performance can only be motivating if the action of employees can influence the measures on which the performance-pay is based. Equally, this evidence does not support those who stress managerial rent-seeking as the principal reason for introducing selective option plans. However in models where endogeneity is accounted for, we do not find any strong evidence of a link with firm productivity. Similarly in models that investigate the dynamics of selective programs, no association between selective options and firm productivity is found. In sum, our findings do not provide strong support for hypotheses of a positive association between option schemes and firm productivity.

Table 1. The evolution of Finnish stock option plans 1992-2002.

Year	(1) # of firms in Helsinki Stock Exchanges	(2) # of new option plans	(3) #of new broad-based option plans	(4) # of firms having option plan	(5) # of firms having broad-based option plan
1992	65	1	0	11 (16.9%)	2 (3.1%)
1993	60	6	1	15 (25.0%)	2 (3.3%)
1994	68	21	2	34 (50%)	3 (5.0%)
1995	74	7	1	38 (51.4%)	3 (4.1%)
1996	73	9	3	36 (49.3%)	6 (8.2%)
1997	115[1)]	22[1)]	4[1)]	46[1)] (40.0%)	7[1)] (6.1%)
1998	119	47	17	69 (58.0%)	21 (17.6%)
1999	137	42	23	91 (66.4%)	36 (26.3%)
2000	150	61	30	113 (75.3%)	54 (36.0%)
2001	145	33	11	112 (77.2%)	54 (37.2%)
2002	137	33	6	101 (73.7%)	49 (35.8%)
Total		282	101		

Notes:
1. Before 1997 data are only for main list firms. From 1997 onwards, the data also include the New Market and the Investor list firms.
2. Note that stock option data in Table 2 also includes firms that have less than four consecutive year observations.
3. Source: Helsinki School of Economics, Alexander Corporate Finance and authors' calculations.

Table 2. The pattern of firm-level panel data, 1992-2002.

Freq.	Percent	Cumulative	Pattern
40	34.2	34.2	11111111111
21	18.0	52.1	01111111111
12	10.3	62.4	00111111111
10	8.6	70.9	00001111111
7	6.0	76.9	00000111111
4	3.4	80.3	00000011111
3	2.6	82.9	00000001111
3	2.6	85.5	00011111111
3	2.6	88.0	01111111110
14	12.0	100.0	(other patterns)
117	100		

Notes
1. The last column describes the pattern of data: 1 means we have an observation for this year, 0 we do not. The first digit (0 or 1) in the pattern column is year 1992. Thus the first row indicates that in 40 cases we have data for all years, whereas the second row indicates that in 21 cases there is no data for 1992 but that data are available in all other years.

Table 3. Summary statistics.

Variable	Name	Firm-year obs	Mean	Std. Dev.
l	Employees	1042	4,066	7,123
k	Fixed capital (tan.+intan.), €1000	1042	464,000	1,500,000
q	Value added, €1000	1042	255,000	574,000
Dilu*	Potential dilution in the range of (0,1); a proxy of option program size	531	0.0547	0.0450
Diluss*	Potential dilution for selective stock option programs	364	0.0286	0.0285
Dilubb*	Potential dilution for broad-based stock option programs	167	0.0900	0.0533
Opt	Option program dummy	1042	0.5182	0.4999
Ssopt	Selective option program dummy	1042	0.3580	0.4796
Bbsopt	Broad-based option program dummy	1042	0.1603	0.3670
Ln(l)	Natural logarithm of employees	1042	7.10	1.62
Ln(k)	Natural logarithm of deflated fixca	1042	17.95	2.07
Ln(q)	Natural logarithm of deflated sales	1042	17.95	1.72

Notes
1. All value measures are deflated using an industry-specific gross output deflator at 2000 constant Euros obtained from Statistics Finland.
2. * Summary statistics for dilu, diluss and dilubb variables are only for those firms that have a stock option program.
3. The total number of firm-year observations is 1042 and data are for 117 firms.

Table 4. Summary statistics: option vs. non-option firms.

Variable	Broad-based option scheme	Selective option scheme	No option scheme
Value added, €1000			
Mean	166,000	496,000	105,000
(Standard deviation)	(366,000)	(834,000)	(238,000)
Employees			
Mean	2,215	7,542	2,100
(Standard deviation)	3,273	9,625	4,371
Fixed capital (tan.+intan.), €1000			
Mean	365,000	929,000	153,000
(Standard deviation)	1,580,000	2,130,000	466,000
Value added / employees, €			
Mean	55,206	57,064	52,191
(Standard deviation)	21,787	18,925	21,302
Firm-year obs.	167	373	502

Notes
1. Based on a firm's option program adoption status in a given year, all firms are classified into three groups, namely broad-based, selective and non-option firms.
2. All value measures are deflated using an industry-specific gross output deflator at 2000 constant Euros obtained from Statistics Finland.

Table 5. Baseline fixed effects estimates: Cobb-Douglas production functions, 1992-2002.

Column	(1)	(2)	(3)	(4)
$\ln(k)_{it}$	0.150 *** (6.38)	0.151 *** (6.39)	0.150 *** (6.38)	0.150 *** (6.37)
$\ln(l)_{it}$	0.617 *** (15.26)	0.621 *** (15.32)	0.615 *** (15.05)	0.619 *** (15.18)
opt_{it}	0.002 (0.06)			
$ssopt_{it}$		0.015 (0.56)		
$bbsopt_{it}$		-0.028 (0.81)		
$dilu_{it}$			0.076 (0.24)	
$diluss_{it}$				0.840 * (1.79)
$dilubb_{it}$				-0.256 (0.73)
Firm-year obs.	925	925	925	925
Firms	117	117	117	117
Baltagi-Wu LBI[1)]	1.32	1.32	1.32	1.32
Modified Bhargava et al.[1)]	0.94	0.94	0.94	0.94
R^2 within	0.61	0.61	0.61	0.62

Notes
1. The dependent variable is ln(value added).
2. The estimator is a modified fixed effects estimator -xtregar-, where the disturbance is first-order autoregressive.
3. Absolute values of t statistics in parentheses. *** Significant at 1% level, ** at 5% level, * at 10% level, respectively.
4. Opt is a dummy variable for the presence of an option program, ssopt is a dummy variable for the presence of a selective program and bbsopt is a dummy variable for the presence of a broad-based option program. Diluss is an interaction variable between potential dilution and ssopt dummy. Dilubb is an interaction variable between potential dilution and bbsopt dummy.
5. Industry-specific year dummies for the ICT, the manufacturing and the service sectors are included in all models.
6. [1)] The tests are based on the standard fixed effects models without modelling first-order autoregression. Baltagi-Wu LBI is the Baltagi-Wu (1999) locally best invariant test statistic for $\rho = 0$. If a test statistic is far below 2, it is an indication of positive serial correlation. Modified Bhargava et al. (1982) also test if $\rho = 0$. If the test statistic is significantly different from zero, we have serial correlation. The tests indicate serial correlation supporting the modified fixed effects estimator.

Table 6. Contemporaneous GMM estimates for 1992-2002 when the option indicator measures the Presence of a program.

Column	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Estimator	OLS level	OLS level	Fixed effects	Fixed effects	Differenced GMM	Differenced GMM	System GMM	System GMM
$\ln(va)_{it-1}$	0.751 *** (17.80)	0.752 *** (17.80)	0.441 *** (5.60)	0.440 *** (5.71)	0.263 *** (3.80)	0.216 *** (2.91)	0.647 *** (8.40)	0.630 *** (9.97)
$\ln(l)_{it}$	0.614 *** (10.30)	0.617 *** (9.96)	0.575 *** (9.11)	0.585 *** (8.81)	0.384 *** (2.76)	0.438 *** (2.84)	0.538 *** (5.43)	0.573 *** (5.94)
$\ln(l)_{it-1}$	-0.430 *** (7.33)	-0.433 *** (7.08)	-0.248 *** (3.08)	-0.252 *** (6.63)	-0.129 (1.42)	-0.125 (1.41)	-0.335 *** (2.89)	-0.368 *** (3.36)
$\ln(k)_{it}$	0.120 *** (4.40)	0.120 *** (4.39)	0.128 *** (4.01)	0.128 *** (4.01)	0.217 *** (3.14)	0.209 *** (2.64)	0.182 *** (4.95)	0.160 *** (3.67)
$\ln(k)_{it-1}$	-0.057 ** (1.98)	-0.057 *** (2.00)	-0.005 (0.16)	-0.005 (0.18)	0.034 (0.78)	-0.030 (0.68)	-0.051 (1.40)	-0.038 (1.23)
opt_{it}	-0.019 (1.00)	-	-0.015 (0.62)	-	-0.053 (0.53)	-	0.044 (0.68)	-
$ssopt_{it}$	-	-0.014 (0.82)	-	0.003 (0.15)	-	0.024 (0.21)	-	0.115 (1.42)
$bbsopt_{it}$	-	-0.031 (0.92)	-	-0.054 (1.15)	-	-0.394 (1.45)	-	-0.097 (1.04)
m1 (p-value)	0.52	0.53	0.93	0.84	-0.01 ***	-0.02 **	-0.00 ***	-0.00 ***
m2 (p-value)	0.23	0.23	0.07 *	0.06 *	-0.34	-0.33	0.77	-0.75
Sargan (p-value)	-	-	-	-	0.21	0.26	0.20	0.34
Firm-year obs.	925	925	925	925	808	808	925	925
Firms	117	117	117	117	117	117	117	117
R^2	0.99	0.99	0.80 (within)	0.80 (within)	-	-	-	-
GMM Instruments	-	-	-	-	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies, Δk_{t-1}, Δl_{t-1}	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies, Δk_{t-1}, Δl_{t-1}

Notes

1. The dependent variable is ln(value added).
2. Absolute values of t statistics are in parentheses, standard errors are heteroskedastic-consistent. *** Significant at 1% level, ** at 5% level, * at 10% level, respectively.
3. GMM estimations are based on the two-step heteroskedastic-robust estimator with a finite-sample correction proposed by Windmeijer (2000).
4. m1- and m2 are tests for first- and second-order autocorrelation in residuals and the statistics are asymptotically standard normal under the null of no serial correlation.
5. Sargan statistic tests the over-identification restrictions. The null hypothesis is that the instruments are valid instruments.
6. Industry-specific year dummies for the ICT, the manufacturing and the service sectors are included in all models.

Table 7. Contemporaneous GMM estimates for 1992-2002 when the option indicator measures the Size of a program.

Column	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Estimator	OLS level	OLS level	Fixed effects	Fixed effects	Differenced GMM	Differenced GMM	System GMM	System GMM	System GMM	System GMM
$\ln(va)_{it-1}$	0.751 *** (17.70)	0.752 *** (17.60)	0.440 *** (5.61)	0.438 *** (5.72)	0.266 *** (3.89)	0.227 *** (3.13)	0.653 *** (8.83)	0.657 *** (8.94)	0.682 *** (10.50)	0.681 *** (10.20)
$\ln(l)_{it}$	0.610 *** (10.20)	0.609 *** (10.30)	0.572 *** (9.03)	0.577 *** (8.97)	0.416 *** (2.48)	0.601 *** (2.75)	0.550 *** (5.17)	0.552 *** (5.32)	0.578 *** (5.85)	0.571 *** (5.48)
$\ln(l)_{it-1}$	-0.428 *** (7.26)	-0.428 *** (7.23)	-0.244 *** (3.10)	-0.242 *** (3.13)	-0.141 (1.48)	-0.115 (1.48)	-0.339 *** (2.97)	-0.341 *** (3.08)	-0.351 *** (3.28)	-0.340 *** (3.86)
$\ln(k)_{it}$	0.119 *** (4.40)	0.119 *** (4.40)	0.128 *** (4.02)	0.128 *** (3.98)	0.236 *** (2.68)	0.201 *** (2.50)	0.181 *** (4.99)	0.179 *** (4.71)	0.152 *** (4.31)	0.149 *** (3.20)
$\ln(k)_{it-1}$	-0.056 ** (1.96)	-0.056 ** (1.96)	-0.001 (0.16)	-0.007 (0.21)	0.027 (0.73)	-0.006 (0.12)	-0.057 (1.60)	-0.060 * (1.72)	-0.063 ** (1.98)	-0.072 ** (1.99)
$dilu_{it}$	-0.059 (0.30)	-	-0.017 (0.05)	-	-1.948 (1.20)	-	0.161 (0.20)	-	-0.381 (0.56)	-
$diluss_{it}$	-	0.019 (0.07)	-	0.728 * (1.72)	-	4.030 (1.07)	-	0.859 (0.55)	-	0.247 (0.29)
$dilubb_{it}$	-	-0.076 (0.34)	-	-0.276 (0.74)	-	-3.462 (1.61)	-	0.061 (0.08)	-	-0.542 (1.07)
m1 (p-value)	0.51	0.51	-0.96	0.89	-0.01 ***	0.01 ***	-0.00 ***	-0.00 ***	-0.00 ***	-0.00 ***
m2 (p-value)	0.22	0.22	-0.07 *	-0.06 *	0.51	-0.59	-0.76	-0.76	-0.79	-0.80
Sargan (p-value)	-	-	-	-	0.26	0.36	0.19	0.21	0.64	0.84
Firm-year obs.	925	925	925	925	808	808	925	925	925	925
Firms	117	117	117	117	117	117	117	117	117	117
R^2	0.99	0.99	0.80 (within)	0.80 (within)	-	-	-	-	-	-
GMM Instruments	-	-	-	-	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies, va_{t-1}, Δk_{t-1}, Δl_{t-1}	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies, Δva_{t-1}, Δk_{t-1}, Δl_{t-1}	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, $dilu_{t-1}$, dummies, Δva_{t-1}, Δk_{t-1}, Δl_{t-1}, $\Delta dilu_{t-1}$	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, $diluss_{t-1}$, $dilubb_{t-1}$, dummies, Δva_{t-1}, Δk_{t-1}, Δl_{t-1}, $\Delta diluss_{t-1}$, $dilubb_{t-1}$

Notes
1. The dependent variable is ln(value added).
2. Absolute values of t statistics are in parentheses, standard errors are heteroskedastic-consistent. *** Significant at 1% level, ** at 5% level, * at 10% level, respectively.
3. GMM estimations are based on the two-step heteroskedastic-robust estimator with a finite-sample correction proposed by Windmeijer (2000).
4. m1- and m2 are tests for first- and second-order autocorrelation in residuals and the statistics are asymptotically standard normal under the null of no serial correlation.
5. Sargan statistic tests the over-identification restrictions. The null hypothesis is that the instruments are valid instruments.
6. Industry-specific year dummies for the ICT, the manufacturing and the service sectors are included in all models.

Table 8. Dynamic GMM estimates for 1992-2002 when the option program indicator measures the Size of a program.

Column	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Estimator	OLS level	OLS level	Fixed effects	Fixed effects	System GMM	System GMM	System GMM	System GMM
$\ln(va)_{it-1}$	0.750 *** (17.60)	0.751 *** (17.40)	0.438 *** (5.68)	0.433 *** (5.77)	0.628 *** (9.09)	0.659 *** (8.47)	0.673 *** (10.50)	0.681 *** (10.20)
$\ln(l)_{it}$	0.609 *** (10.20)	0.608 *** (10.20)	0.573 *** (8.99)	0.580 *** (8.91)	0.555 *** (5.13)	0.555 *** (4.90)	0.584 *** (6.04)	0.557 *** (6.36)
$\ln(l)_{it-1}$	-0.425 *** (7.18)	-0.426 *** (7.12)	-0.240 *** (3.13)	-0.235 *** (3.16)	-0.323 *** (2.98)	-0.338 *** (3.00)	-0.343 *** (3.42)	-0.326 *** (3.41)
$\ln(k)_{it}$	0.119 *** (4.42)	0.119 *** (4.42)	0.129 *** (4.02)	0.128 *** (3.98)	0.179 *** (4.45)	0.173 *** (3.87)	0.150 *** (4.04)	0.147 *** (4.03)
$\ln(k)_{it-1}$	-0.056 ** (1.96)	-0.056 ** (1.97)	-0.005 (0.15)	-0.006 (0.20)	-0.051 (1.43)	-0.062 * (1.72)	-0.061 ** (1.98)	-0.066 *** (2.06)
$dilu_{it}$	0.209 (0.74)	-	0.177 (0.59)	-	1.093 (0.20)	-	0.089 (0.19)	-
$dilu_{it-1}$	-0.340 (0.94)	-	-0.310 (0.87)	-	-1.536 (0.96)	-	-0.555 (1.12)	-
$diluss_{it}$	-	0.407 (0.77)	-	0.500 (1.04)	-	2.800 (1.00)	-	1.192 (0.77)
$diluss_{it-1}$	-	-0.540 (0.85)	-	0.314 (0.71)	-	-3.424 (1.20)	-	-1.610 (1.09)
$dilubb_{it}$	-	0.106 (0.38)	-	0.043 (0.13)	-	0.425 (0.32)	-	-0.016 (0.03)
$dilubb_{it-1}$	-	-0.229 (0.60)	-	-0.432 (1.01)	-	-0.726 (0.44)	-	-0.576 (0.78)
Wald test (p-value)	0.64 (dilu)	0.69 (diluss)	0.64 (dilu)	0.24 (diluss)	0.63 (dilu)	0.49 (diluss)	0.53 (dilu)	0.55 (diluss)
Wald test (p-value)	-	0.83 (dilubb)	-	0.59 (dilubb)	-	0.91 (dilubb)		0.72 (dilubb)
m1 (p-value)	0.51	0.51	0.96	0.93	-0.00 ***	-0.00 ***	-0.00 ***	-0.00 ***
m2 (p-value)	0.23	0.23	-0.07 *	-0.06 *	-0.73	-0.80	-0.77	-0.78
Sargan (p-value)	-	-	-	-	0.17	0.17	0.60	0.95
R^2	0.99	0.99	0.80 (within)	0.80 (within)	-	-	-	-
GMM Instruments	-	-	-	-	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies, Δva_{t-1}, Δk_{t-1}, Δl_{t-1}	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, dummies, Δva_{t-1}, Δk_{t-1}, Δl_{t-1}	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, $dilu_{t-1}$, dummies, Δva_{t-1}, Δk_{t-1}, Δl_{t-1}, $\Delta dilu_{t-1}$	va_{t-2}, va_{t-3}, k_{t-1}, k_{t-2}, l_{t-1}, l_{t-2}, $diluss_{t-1}$, $dilubb_{t-1}$, dummies, Δva_{t-1}, Δk_{t-1}, Δl_{t-1}, $\Delta diluss_{t-1}$, $\Delta dilubb_{t-1}$

Notes
1. The dependent variable is ln(value added).
2. Absolute values of t statistics are in parentheses, standard errors are heteroskedastic-consistent. *** Significant at 1% level, ** at 5% level, * at 10% level, respectively.
3. GMM estimations are based on the two-step heteroskedastic-robust estimator with a finite-sample correction proposed by Windmeijer (2000).
4. m1- and m2 are tests for first- and second-order autocorrelation in residuals and the statistics are asymptotically standard normal under the null of no serial correlation.
5. The Wald test tests whether contemporaneous and lagged dilution option indicators are jointly statistically significant.
6. Sargan statistic tests the over-identification restrictions. The null hypothesis is that the instruments are valid instruments.
7. Industry-specific year dummies for the ICT, the manufacturing and the service sectors are included in all models.
8. The number of firms is 117 or 925 firm-year observations in all models.

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Endnotes

[1] Earlier versions of this paper have benefited from comments by participants at the ASSA/ACES meeting in San Diego, January 2-5, 2004, the 12th IAFEP Conference in Halifax, July 8-10, 2004, the 16th EALE conference in Lisbon, September 9-11, 2004, the FPPE Industrial Organisation Workshop in Helsinki, December 9-10, 2004, and the EALE/SOLE World Conference in San Francisco, June 2-5, 2005. We are especially grateful to Kari Hämäläinen, Pekka Ilmakunnas, Uwe Jirjahn, Jeffrey Pliskin and Otto Toivanen for their very helpful comments. Also we acknowledge Professor Seppo Ikäheimo and Alexander Corporate Finance for allowing us access to their databases on options in publicly traded Finnish companies, and to Mikael Katajamäki for his outstanding research assistance. We also thank Balance Consulting for financial statement data. Kalmi and Mäkinen gratefully acknowledge funding from the LIIKE-programme of the Academy of Finland. Kalmi also gratefully acknowledges financial support from the Marcus Wallenberg Foundation and the Helsinki School of Economics Research Foundation. In addition, Mäkinen thanks the Yrjö Jahnsson Foundation, the Helsinki School of Economics Research Foundation and the Foundation of Kluuvi for financial support. Support from the Research Institute of the Finnish Economy (ETLA) is gratefully acknowledged.

[2] Mäkinen (2001) describes the evolution of stock option programs in Finland. Jones et al. (2006) study the determinants of option schemes adoption in Finland. They also summarise the evolution of options in Finland and discuss the institutional background in more detail.

[3] These are part of a broader class of studies that employ an augmented production function methodology. For example, for a review of such work in investigating the impact of ownership forms on firm performance in transition economies, see Djankov and Murrell (2002).

[4] However, we have included option schemes prior to the listing for such firms that enter the HEX before 2002.

[5] This is done to provide the reader with a better understanding of the development and prevalence of option schemes in Finland. See more detailed evolution and institutional background description in Jones et al. (2006).

[6] We omit 8 firms or 15 firm-year observations due to their having fewer than 4 consecutive observations. To utilize all possible firm-level financial information data we also collected data on income statements prior a firm's listing on the Helsinki Stock Exchange.

[7] Our classification is thus different from Kroumova et al. (2000, 2002), who use 50 % threshold as a criterion for broad-based schemes. Our data do not include this information, but have the important advantage of being derived from publicly reported sources that must be externally verifiable, rather than from confidential surveys.

[8] We also interviewed Mr. Erkki Helaniemi, a partner in the investment bank Alexander Corporate Finance, who has been personally involved in setting up dozens of option schemes. He confirmed that there are dramatic differences in the participation rates for option schemes, depending on eligibility.

[9] The first two are option program dummy variables, which measure the presence or absence of an option program. The dummy variable captures a program's "introduction effect": when an option program is adopted, a dummy variable switches from 0 to 1 shifting a linear production function immediately and without anticipation. This effect of an option plan on firm-level productivity is modelled as a constant over firms. The third measure is a proxy for the size of a program.

[10] Since we do not have access to information on stock option program details such as exercise prices, we cannot calculate Black-Scholes values. However, Hallock and Olson (2006) find that employees may value their stock options significantly above the Black-Scholes value. See also Bergman and Jenter (2006).

[11] When instrumenting inputs used in a production process in a dynamic GMM estimation, the instrument matrix becomes substantially larger in the translog case. If the full set of instruments is used, this may bias estimates.

[12] For example, presumably the terms of option schemes differ among firms. Thus, when an option scheme is substantially out of the money (i.e. the current stock price is substantially below the exercise price), options may not provide strong incentives for employees and managers to improve their performance.

[13] On theoretical grounds firm value added is a preferable measure to sales, since value added does not include intermediate inputs that are purchased from other firms.

[14] This decision is supported by the fact that the ICT sector experienced both boom and bust during 1997-2003.

[15] Our decision to use a GMM estimator has been guided by the fact that we do not have suitable instrument variables for option schemes.

[16] Besides these dynamic panel data GMM estimators, Maximum likelihood (ML) estimators have also been developed. Unfortunately, one drawback of the ML estimators is that different distributional assumptions of the initial conditions are needed in the estimation process to imply different likelihood functions. It follows that ML estimators may produce inconsistent estimate for the lagged dependent variable (i.e. $\ln va_{,t-1}$), if an initial condition distribution, i.e. the distribution of va_{i1}, is not the correct one. By contrast, the dynamic panel data GMM estimators do not need such a strong initial condition assumption.

[17] This seems to be a reasonable assumption, since firms may not adjust their capital and labour inputs immediately on economic shocks within a year. For example, firms may be unaware as to whether a shock is permanent or temporary.

[18] Since option plans typically are introduced publicly in early spring, a few weeks before the annual general meeting of shareholders, this seems to be a reasonable assumption. Therefore, a potential correlation is likely to be with a previous period rather than being contemporaneous. We also estimated an IV estimator (i.e. -ivreg2- with -cluster- in the Stata), where an option program indicator was treated as an endogenous variable in a Cobb-Douglas production function. As excluded instruments we used the HEX general index, ICT sector dummy indicator and the share ownership of a largest shareholder. The instrumented option dummy indicator was insignificant (p-value 0.84), when the F-test for the excluded instruments in the first-stage regression was highly significant (p-value 0.00) and the Hansen J statistic supported the validity of the (all) instruments (p-value 0.12).

[19] For more on dynamic GMM estimators and constructing an instrument variable matrix see Arellano and Bond (1991), Arellano and Bover (1995), and Blundell and Bond (1998).

[20] The benefit of using the system GMM estimator is that it is shown to be more efficient than the differenced GMM estimator in large samples.

[21] To us the non-correlation seems more plausible approach than assuming that the first-differences of capital and labour are uncorrelated with individual effects η_i.

[22] We use Stata/SE 9.1 for Windows statistical package in estimating the models of Table 5.

[23] We also performed the modified Wald statistics to examine a groupwise heteroskedasticity in the residuals when estimating a standard fixed effects estimator. The statistics indicate a violation in assumption that the error term is homoskedastic. Thus, the modified Wald statistic indicates that the fixed effects estimator is not efficient, which challenges our conclusions on the significance levels of the standard fixed effects parameters. However, the modified Wald statistics should be also used with caution, since its power can be low in the context of fixed effects with "large N, small T". Unfortunately, we are unaware of any fixed effects estimators that account simultaneously for autocorrelation and heteroskedasticity. We estimate heteroskedastic robust estimates in the GMM models; therefore we prefer accounting for autocorrelation in Table 5. The autocorrelation ρ is calculated by -tscorr- option in -xtregar- model by using the Stata/SE 9.1 for Windows statistical package. The benefit of using -tscorr- is that ρ is bounded in [-1,1].

[24] While these estimates would imply decreasing returns to scale, it is well known that fixed effects estimators tend to underestimate the coefficients, especially capital (Griliches and Mairesse, 1998).

[25] All estimations use the Ox/DPD statistical package. This econometric software package allows tests for the first- and second-order autocorrelation and the Sargan test for over-identifying restrictions. It also calculates asymptotically heteroskedastic robust standard errors. For more on Ox/DPD statistical package see http://www.doornik.com.

[26] See e.g. Bond (2001) and Arellano (2003) for technical details on a two-step GMM estimator.

[27] We prefer a two-step GMM estimator for two reasons. First, it should be more efficient in the presence of heteroskedasticity, especially under Windmeijer's (2000) finite sample correction. Second, the Sargan statistic based on the minimised value of the two-step GMM estimator has an asymptotic χ^2 distribution regardless of heteroskedasticity.

[28] The estimation results suggest (not reported here) that the individual series are highly persistent (but not an exact unit root) indicating that the instruments in first-differenced equations are likely to be weak. Being the case, the existing evidence from standard instrumental variable literature suggests that IV estimators can be subject to serious finite sample biases when instruments are weak (e.g., Bound et al. (1995)). Blundell and Bond (1998), and Blundell et al. (2000) have demonstrated that this also applies for

the differenced GMM estimator, when individual series are highly persistent biasing the estimates downwards.

[29] The information on β's in the first-difference regressions depend on the ratio of var (Δv)/var (Δx). Variation of x over time t is necessary, as is variation over individuals i. Since in our data the i,t -variation of option program dummy indicators is less than variation in the dilution indicators, we focus on the correlation between the dilution option program indicator and the error term.

[30] As in Table 6 previously, also now the differenced GMM estimates for va_{t-1} in columns (5) and (6) are clearly below the fixed effects estimates in columns (3) and (4). This again indicates a serious finite sample bias for the estimator, which is likely to be associated with weak instruments. Moreover, the autocorrelation tests m1 and m2 indicate that the key assumption for the system GMM estimators is fulfilled. Also, the system GMM estimates for va_{t-1} in columns (7)-(10) are lower than the OLS level estimates but higher than the fixed effects estimates indicating that the system GMM estimator is likely to be consistent, at least for lagged dependent variable.

[31] As in Tables 6 and 7, the differenced GMM estimators appeared to be inconsistent in Table 8. Consequently, we do not report the estimates for the differenced GMM estimator in Table 8.

Do Stock Option Schemes Affect Firm Technical Inefficiency? Evidence from Finland[1]

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Abstract:
In this paper we study whether stock option schemes affect firm technical inefficiency. We estimate Cobb-Douglas stochastic production frontier models using a novel panel data set on the publicly listed Finnish firms in the manufacturing and ICT sectors over the period from 1992 to 2002. We find evidence that the mean inefficiency estimates in the ICT sector are clearly higher than in the manufacturing sector. Furthermore, our empirical findings suggest that broad-based option firms may have higher mean inefficiency than selective and non-option firms in the manufacturing sector. The quantitative assessments of the marginal effects on the inefficiency support the view that especially broad-based schemes affect the mean and the variance of the inefficiency term u_{it} in the manufacturing sector, but not in the ICT sector. Our findings do not provide empirical support for the view that stock option schemes reduce firm technical inefficiency.

Keywords: stochastic frontier, technical inefficiency, production function, stock options

JEL-codes: C3, J3, M5

1. Introduction

During the 1990s, stock options became an increasingly popular compensation method in many countries (e.g. Murphy, 1999). Initially, stock options were typically allocated only to executives[2], but the association of stock options mainly with managerial compensation changed rapidly after companies worldwide started to issue options to the workforce more broadly (e.g. Weeden, Carberry and Rodrick, 1998; Lebow et al., 1998; and Blasi, Kruse and Bernstein, 2003). The growing use of stock options has generated heated public discussion with some viewing stock options as a device by which managers transfer excessive benefits to themselves, while others see options as a major innovation in managerial and personnel compensation.

The growth of option adoptions has accompanied a mushrooming of theoretical and empirical literature on stock options (e.g. Ittner et al., 2003). Whereas sharp disagreements exist among theorists on the economic impact of different types of option schemes, an existing empirical work in economics has typically focused on the link between options and firm productivity. For example, Jones, Kalmi and Mäkinen (2006b) argue: *"For selective option schemes, the baseline fixed effects estimator suggests a 2.1-2.4% positive and statistically significant effect of the option program indicator on firm productivity. However, in empirical models in which endogeneity and dynamics are taken into account, no evidence is found of a link with firm productivity."* This evidence of a non-significant link raises a question whether, instead of firm productivity, stock options affect firm technical inefficiency, as inefficiency is defined in the stochastic production frontier literature. For example, the proponents of options typically argue that option plans may motivate managers and employees to make better decisions, work harder and share information within a firm in a way that decreases firm inefficiency. Other examples of exogenous factors that may affect inefficiency are the degree of competitive pressures, input and output quality indicators, network characteristics, ownership form, and various managerial characteristics etc. (Kumbhakar and Lovell, 2000).[3] To the best of our knowledge, we provide the first empirical evidence in the literature on the link between stock option schemes and firm technical inefficiency.

The key research questions are: (i) whether firm-level technical inefficiency is higher in non-option than in option firms; (ii) whether the impact of options on firm technical inefficiency is dependent upon whether a plan is broad-based or selective. We estimate simultaneously stochastic production frontier parameters, inefficiency scores and marginal effects by using novel panel data on Finnish publicly listed firms[4] in the manufacturing and ICT sectors in 1992-2002. Our data enable a careful investigation of the inefficiency effects of different types of option plans, i.e. whether options are allocated selectively to a specific group of employees (i.e. a selective option scheme) or whether all employees are eligible to participate (i.e. a broad-based option scheme). Since a possibility to obtain firm-level inefficiency estimates is the main reason to use stochastic frontier models, we follow a common procedure in the literature and treat all explanatory variables as exogenous.

We find evidence that the shape of the inefficiency distribution differ notably between the manufacturing and the ICT sectors. For example, mean inefficiency estimates in the ICT sector are substantially higher than in the manufacturing sector, though naturally efficient and inefficient firms exist in both sectors. Also, in the ICT sector mean conditional inefficiency estimates indicate that there is no mean inefficiency difference between option and non-option firms. However, in the manufacturing sector our findings suggest that broad-based firms may have higher mean inefficiency than selective and non-option firms.

The quantitative assessment of the average marginal effects on the inefficiency term supports the view that especially broad-based schemes affect the mean and the variance of the inefficiency term in the manufacturing sector. The findings on the mean of inefficiency suggest that broad-based schemes may increase technical inefficiency. Respectively, the marginal effect of broad-based schemes on the variance of the inefficiency term is significant, implying an increase in production uncertainty. In sum, these findings would indicate, other things equal, that broad-based scheme firms in the manufacturing sector may achieve lower and more uncertain productivity growth as time goes by. For selective schemes, we find no evidence of a link with technical inefficiency. Finally, our findings do not support the hypothesis that option schemes reduce firm technical inefficiency.

This paper is organised as follows. Section 2 briefly describes the evolution of stock option programs in Finland. In section 3 we describe our data and empirical strategy. Section 4 reports the empirical findings. Finally, section 5 concludes.

2. The development of option schemes in Finland

In this section, we briefly review the option schemes' adoption pattern in Finland.[5] Table 1 describes the evolution of option plans in the publicly traded firms on the Helsinki Stock Exchange (HEX) between 1987 (when the first employee stock option scheme was launched in Finland) and 2002. We have information on the presence of option schemes on the main list throughout the period and on the minor lists, i.e. NM-list (New Market) and I-list (Investor), since 1997.

Column 1 gives the number of firms on the HEX main list. Column 2 shows the total number of listed firms, including the two minor lists (from 1997). It appears that the number of listed firms fluctuates a lot with the business cycle. The first period of growth was the economic boom years 1987-1989, when the number of firms increased from 52 to 82. From 1989 onwards the number of firms fell, reaching a low point of 60 firms in 1993. The main reason for this was the Great Finnish Depression in 1990-1993, when many Finnish firms had financial problems.[6] After 1993 the number of listed firms started to rise, and the 1989 level was reached again in 1997. The increase continued until 2000, but thereafter the number fell again. From 1997 onwards we also include firms on the two minor lists. In some cases, firms switched from the minor to the major list. At the same time, however, there are new firms entering the minor lists, especially in 2000 when relatively many small ICT firms entered the NM-list.

Column 3 indicates how many firms have adopted their first option scheme in a given year. Altogether, 127 firms have adopted a stock option plan. While seven pioneering firms implemented their option plans as early as the 1980s, very few plans were launched during the economic depression years of 1990-1993. The renewed interest in option plans began in 1994, when 20 firms (almost 40% of listed firms) adopted option schemes. Relatively few firms adopted schemes during 1995-1996 (possibly because the taxation of option gains changed from a moderate capital tax into a substantially higher marginal income tax), but since 1997 options have became widely popular. The rise of option schemes during 1999-2000 was fuelled by new listings. When new listings stopped after 2000, so did the introduction of new option schemes. Firms often launch new schemes once the previous schemes are close to expiring, or they may operate many schemes simultaneously: 84 of the 127 firms (66%) that have ever adopted a scheme have implemented more than one scheme (three firms have reached 7 successive schemes).[7]

Column 4 shows the number of firms that adopted new option schemes in a given year. The total number of option adoptions we are aware of is 290. The early peak year was 1994 (21 firms adopted). From 1997 (22 firms adopted) the adoption increased further, but after 59 plans in 2000 the adoptions started to decline, with 32 new schemes in 2001 and 28 in 2002.

In Column 5 we use the information on timing and launching of a scheme. A firm is treated as having a scheme in year t, if it has at least one scheme that has started in year t or earlier and if the final date for exercising options in this scheme is in year t+1 or later. Column 5 indicates that the proportion of firms with an option scheme increased until 1993, by which time 20% of the main list firms had an option scheme. This proportion jumped to around 40% in 1994, after which it increased slowly for three years, until it jumped again to 65% in 1998. The temporary maximum was reached in 2001, when almost 85% of the main list firms had a stock option scheme.

Column 6 shows the development for all firms, also for those outside the main list. The proportion of firms with stock option schemes is somewhat lower for all firms, due to many non-option firms at the I-list.

More generally, the extensive growth of stock option schemes reflects a deep change in the Finnish corporate governance system. In the end of the 1980s, the Finnish corporate governance system in listed firms was very much bank-centred and resembled the German system (see e.g. Hyytinen, Kuosa and Takalo, 2003). The stock market started its recovery after the depression in 1993, and the importance of the equity market in financial intermediation grew throughout the 1990s. Both the turnover and market value of firms listed on the stock exchange increased dramatically throughout the decade, with Nokia leading this development.

Now Finnish stock markets are much deeper, more transparent and arguably provide more reliable information than in the past. At the same time, both monitoring of insider trading and legal punishments have become stricter. During the last 10-15 years Finland has shifted from a bank-based financial intermediation towards a market-based system. As discussed above, the most active period of stock option adoptions coincided with the height of the stock market boom in the late 1990s. However, as market prices started to fall after May 2000, accelerating further in 2001 and 2002, the rate of stock option adoption decreased markedly (see e.g. Jones, Kalmi and Mäkinen, 2006a).

3. The data and estimation strategy

3.1 The data

In this section we describe the data and our estimation strategy. To examine the impact of option schemes on firm technical inefficiency, we use new panel data for the publicly listed Finnish firms in the manufacturing and ICT sectors in 1992-2002. Our firm-level data include information on firms' stock option programs and financial statements. Moreover, our option data enable a distinction between selective and broad-based schemes allowing an investigation of the inefficiency effects of different types of option plans. In the option data set, we have combined four different option data sources: firms' annual statements and general meeting reports, firms' press releases on the adoption of a scheme, the option data gathered by Professor Seppo Ikäheimo from the Helsinki School of Economics, and the option data provided by Alexander Corporate Finance Ltd, an investment bank that designs option programs in Finland. We then cross-checked the option information several times, and in a few cases when it did not match, we have trusted the firms' own public announcements. Thereafter we matched option data with firm-level accounting data, obtained from Balance Consulting Ltd, a firm specialised in accounting information.

Our data include *all* listed Finnish companies for a minimum of four consecutive years in the manufacturing and ICT sectors. It is an unbalanced panel, i.e. we do not observe the same cross-section units in each year. Apparently, some of the yearly variation is due to the entry and exit (attrition) of listed firms at the Helsinki Stock Exchange (HEX). Also, a few firms merged in the period, and in these cases we included only new merged firms after the merger. In addition, concerning mainly recently listed firms in a few cases, we added a firm's financial statement information prior to the listing, if that information was available in the accounting data.[8]

To control for potential bias of very small and very large firms, we have excluded potential outlier observations, i.e. an observation if employment was less than 50 persons, if fixed capital was less than €1,000,000, and if employment was more than 50,000 persons. We also deflated all nominal monetary variables by an industry based gross-output deflator at constant 2000 Euros, obtained from Statistics Finland. The final data set contains 571 firm-year observations regarding 62 firms in the manufacturing sector and 243 firm-year observations covering 32 firms in the ICT sector, so that the number of observations of a firm *i*, i.e. T_i, is $4 \leq T_i \leq 11$.[9] Tables 2 and 3 present summary statistics for our key variables in the manufacturing and the ICT sectors.

[Tables 2 and 3 about here]

In the analysis that follows, we distinguish between broad-based and selective option schemes. The latter are mostly managerial schemes, although they can also include other key personnel (e.g. R&D employees). However, in order to qualify as a broad-based scheme, all employees (or at least a great majority) should be eligible to participate. The classification on broad-based and selective option schemes is based on firms' public stock exchange reports.[10] The Finnish Law on Joint Stock Companies requires listed firms to report all relevant terms of stock option schemes to shareholders prior to adoption. While a high rate of eligibility does not automatically guarantee a high participation rate, there are good reasons to believe that these are closely connected. For one thing, employees usually face only small costs when they subscribe to options—e.g. by providing a zero-interest loan to the company, with the company repaying the loan at face value after a certain period, usually 1-3 years. Thus, while employees face a cost in terms of foregone interest and liquidity, typically this cost is far below the real value of the options. Moreover, not all companies use this procedure, as they essentially give options for free to their employees.[11]

We use different indicators for the presence or absence of an option scheme, the size of the scheme and whether the scheme is selective or broad-based. Two of the indicators are binary variables and one is a continuous variable. Our *first binary indicator* is *opt* measuring the presence of a scheme in a firm in a given year t. It equals one for option firms and zero otherwise. Thus, the indicator distinguishes option and non-option firms and allows us to compare inefficiency differences between option and non-option firms.

Our *second binary indicator* measures also the presence or absence of a plan, but it distinguishes between selective *(ssopt)* and broad-based *(bbsopt)* plans. By a selective plan we mean a scheme that is targeted to a selected group of employees including managerial programs, but also schemes that are targeted to key personnel. Broad-based plans are all encompassing, including managers, but they do not have to be egalitarian in the sense of all participants having the same number of options. By using these distinct dummy variables, we can examine whether inefficiency differs between selective and broad-based schemes.

Our *third program indicator* is potential dilution *(dilu)*. This indicator measures the potential size of effective schemes in a firm in a given year.[12] This is a continuous variable, i.e. the ratio of the number of shares that may be awarded through effective option plans in a given year divided by the sum of the total number of shares and the number of new shares that may be awarded through options at the end of a year. If a program ends in the middle of a year t, then year t-1 is the last year used in assessing a potential dilution. The indicator allows us to explore whether option schemes can be simultaneously associated with the mean and the variance of the inefficiency term. We also use dilution indicators for selective (*diluss*) and broad-based (*dilubb*) schemes to examine whether there is a difference between the schemes.

3.2 Estimation strategy

In their pioneering work Aigner, Lovell and Schmidt (1977) and Meeusen and van den Broeck (1977) proposed independently stochastic production frontier models. Since then the literature has proposed several specifications and estimation techniques.[13] Early specifications focused on estimating technical inefficiency with cross-section data, but access to panel data allowed a richer modelling approach in the form of the fixed effects (e.g. Schmidt and Sickles, 1984) and the random effects (e.g. Pitt and Lee, 1981) estimators enabling us to relax some relatively strong distributional assumptions needed in cross-section models.[14] A stochastic production frontier panel data model can be written as

(1) $y_{it} = \alpha + \beta' x_{it} + v_{it} - u_{it}$, $u_{it} \geq 0$, i=1,...,N and t=1,...,T,

where firm output y_{it} is a scalar, and x_{it} is a vector of explanatory variables, such as inputs used in a production process. As proposed by Aigner, Lovell and Schmidt (1977) and Meeusen and van den Broeck (1977), a composed error term $\varepsilon_{it} = v_{it} - u_{it}$ is the difference between a normally distributed two-sided noise component v_{it} and a normally distributed nonnegative inefficiency component u_{it} with the following assumptions:[15]

(2) $v_{it} \sim \mathrm{N}\left(0, \sigma_v^2\right)$, $u_{it} \sim \left|U_{it}\right|$ where $U_{it} \sim \mathrm{N}\left(0, \sigma_u^2\right)$ and independent of v_{it}.

A research interest may be in production technology parameters β, but one of the main reasons to estimate stochastic frontier models lies in obtaining inefficiency estimates $\hat{u}_{it}$. Unfortunately, these cannot be obtained directly from Equation (1), since only composed residuals $\hat{\varepsilon}_{it}$ are observed. Jondrow, Materov, Lovell, and Schmidt (1982) proposed the mean of the conditional distribution of u_{it} given ε_{it} as a point estimator for inefficiency term under the distributional assumptions presented in Equation (2):

(3) $$\mathrm{E}\left[u_{it} \middle| \varepsilon_{it}\right] = \frac{\sigma\lambda}{1+\lambda^2}\left[\frac{\phi\left(\frac{\varepsilon_{it}\lambda}{\sigma}\right)}{1-\Phi\left(\frac{\varepsilon_{it}\lambda}{\sigma}\right)} - \left(\frac{\varepsilon_{it}\lambda}{\sigma}\right)\right],$$ where

$\sigma = \left(\sigma_u^2 + \sigma_v^2\right)^{1/2}, \lambda = \frac{\sigma_u}{\sigma_v}$, and $\Phi(\cdot)$ and $\phi(\cdot)$ are standard normal cumulative distribution and density functions, respectively. Although Equation (3) gives point estimates for conditional technical inefficiencies[16], a major drawback is that it does not assess what may drive these inefficiencies.[17]

Greene (2002a; 2005) and Wang (2002) have recently proposed new maximum likelihood estimators that provide parameterizations of the exogenous influences on inefficiency. For example, Greene (2002a; 2005) suggests several estimators to account for heterogeneity among firms and to estimate simultaneously both technology parameters and technical inefficiency.[18] Wang (2002) proposes a model where heteroscedasticity and non-monotonic efficiency effects can be modelled. In addition, the model allows one to accommodate unconditional marginal effects of exogenous variables on the mean and the variance of u_{it} and to examine statistical significance of marginal effects by bootstrapping.[19]

We denote a firm's production function by $f(.)$, which relates firm value added[20] at time t, i.e. va_{it}, to inputs used in a production process:

(4) $\quad va_{it} = f(k_{it}, l_{it}, x_t; \beta)$, where i=1,2, ..., N and t=1,2,...,T.

In Equation (4), subscripts i and t index firm and time, respectively. Firm deflated fixed capital is k_{it}, the sum of a firm's tangible and intangible assets at the end of the year, labour input l_{it} is the mean number of employees in a given year, and x_t is a time trend to account for technological change. We assume a Cobb-Douglas stochastic production frontier[21] as follows:

$$\begin{aligned} &\ln va_{it} = \beta_k \ln k_{it} + \beta_l \ln l_{it} + \beta_x x_t + \varepsilon_{it}, \text{ where i=1,2,..., N; t=1,...,T,} \\ (5) \quad &\varepsilon_{it} = v_{it} - u_{it}, v_{it} \sim \mathrm{N}\left(0,\sigma_v^2\right),\ u_{it} \sim \left|U_{it}\right|,\ U_{it} \sim \mathrm{N}\left(0,\sigma_u^2\right). \end{aligned}$$

In Equation (5) all variables are the same as in Equation (4): a firm's inputs in the production process are capital k_{it} and labour l_{it}, and x_t is a linear time trend. We estimate separate industry-level models for the ICT and manufacturing sectors, since the sectors may differ in several ways. For example, a firm may need more capital in the manufacturing sector than in the ICT sector, whereas labour may be a more important production factor in the ICT than in manufacturing.

To study whether stock option schemes affect firm technical inefficiency, we utilize the recent developments in the literature that allows parameterizing the composite error term $\varepsilon_{it} = v_{it} - u_{it}$.[22] By doing this we can account for heteroscedasticity in the inefficiency component u_{it} (e.g. Caudill and Ford, 1993) and in the noise component v_{it} (e.g. Hadri, 1999).[23] But more importantly, by modelling the mean and the variance of inefficiency term u_{it}[24] as a function of stock option schemes, we can examine our two research hypotheses, namely (i) that firm-level technical inefficiency is expected to be higher in non-option than option firms; (ii) that the impact of options on firm technical inefficiency is expected to be dependent upon whether the plan is broad-based or selective. Thus, the variance of u_{it} is parameterized as an exponential function of firm size (measured by $\ln(l_{it})$) and stock option variables as follows[25]:

$$(6) \qquad \sigma_u^2 = \sigma_{uit}^2 = \exp(\delta' z_{it}) = \exp(\alpha + \delta_L \ln(l_{it}) + \delta_{opt} opt_{it})$$

$$(7) \qquad \sigma_u^2 = \sigma_{uit}^2 = \exp(\delta' z_{it}) = \exp(\alpha + \delta_L \ln(l_{it}) + \delta_{ssopt} ssopt_{it} + \delta_{bbsopt} bbsopt_{it}).$$

Besides the variance of the inefficiency component, the symmetric noise component can be heteroscedastic with respect to the size of firms. Thus, we model v_{it} as an exponential function of firm size as follows:

(8) $$\sigma_v^2 = \sigma_{vit}^2 = \exp(\gamma' z_{it}) = \exp(\alpha + \gamma_L \ln(l_{it})).$$

To model flexible parameterizations of exogenous influences on the mean (e.g. Kumbhakar, Ghosh and McGuckin, 1991) and the variance of the inefficiency term u_{it} (e.g. Caudill and Ford, 1993), we use a model suggested by Wang (2002).[26] Contrary to Equations (6)-(8), now the effects on the inefficiency term are measured by the unconditional statistics of $\mathrm{E}[u_{it}]$ and $\mathrm{Var}[u_{it}]$.[27] The first two moments of u_{it} are

(9) $$m_1 = \mathrm{E}[u_{it}] = \sigma_{it}\left[\Lambda + \frac{\phi(\Lambda)}{\Phi(\Lambda)}\right]$$

(10) $$m_2 = \mathrm{Var}[u_{it}] = \sigma_{it}^2\left[1 - \Lambda\left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right] - \left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right]^2\right], \text{ where}$$

$\Lambda = \dfrac{\mu_{it}}{\sigma_{it}}$, and $\Phi(\cdot)$ and $\phi(\cdot)$ are standard normal cumulative distribution and density functions, respectively. The marginal effect of an exogenous variable z on $\mathrm{E}[u_{it}]$ can be calculated as follows:

(11) $$\frac{\partial \mathrm{E}[u_{it}]}{\partial \mathrm{z}} = \delta_z\left[1 - \Lambda\left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right] - \left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right]^2\right]$$

$$+\gamma_z \frac{\sigma_{it}}{2}\left[(1+\Lambda)^2\left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right] + \Lambda\left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right]^2\right], \text{ where}$$

δ_z and γ_z are the estimated coefficients of an exogenous variable z in Equations (6)-(8). Thus, the marginal effect is the sum of adjusted slope coefficients. Respectively, the marginal effect of an exogenous variable z on $\mathrm{Var}[u_{it}]$ is

$$(12)\quad \frac{\partial \mathrm{Var}[u_{it}]}{\partial z} = \frac{\delta_z}{\sigma_{it}}\left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right](m_1^2 - m_2) +$$

$$+\gamma_z \sigma_{it}^2\left[1 - \frac{1}{2}\left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right]\left(\Lambda + \Lambda^3 + \left(2+3\Lambda^2\right)\left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right] + 2\Lambda\left[\frac{\phi(\Lambda)}{\Phi(\Lambda)}\right]^2\right)\right], \text{ where}$$

m_1 and m_2 are the first two moments of u_{it}, represented in (9) and (10).[28]

4. Estimation results

Table 4 presents the stochastic production frontier estimates for the manufacturing and ICT sectors. As can be seen from Table 4, production technology parameter estimates are in line with our prior expectations, i.e. capital input elasticities are higher in manufacturing than in the ICT sector, whereas labour input elasticity estimates are higher in the ICT sector than in manufacturing. In the manufacturing sector, estimated elasticities for capital are 0.29 (0.16 in the ICT sector) and for labour 0.71 (0.85 in the ICT sector), indicating that a production process in the manufacturing sector is more capital and less labour intensive than in the ICT sector. In both sectors, Wald tests support constant returns to scale hypothesis. The constant rate of technical change estimate is about 2.5 % per year in the manufacturing sector, but we find no evidence of that in the ICT sector. The estimates of λ are statistically significant and higher than one, indicating the existence of inefficiency in both sectors. When comparing the estimates of σ_u between the sectors, the variance of technical production inefficiency appears to be clearly higher in the ICT sector (0.68) than in the manufacturing sector (0.18). In addition, especially the presence of broad-based schemes seems to affect the variance of σ_u. The choice of parameterizing the error term v_{it} as a function of firm size seems to be an adequate approach in the manufacturing sector, but the size of the firm is statistically insignificant in the ICT sector.

Based on stochastic production frontier models in Table 4, Tables 5-6 report conditional inefficiency estimates by industry and the type of option scheme. In Table 5, the mean inefficiency estimates are substantially higher in the ICT sector than in manufacturing. For example, the estimated mean inefficiency is 0.21 or 21% with a standard deviation of 0.13 in the manufacturing sector, whereas in the ICT sector it is 0.45 or 45% with a standard deviation of 0.26. It is, however, important to notice that efficient and inefficient firms exist in both sectors, e.g. minimum inefficiency is 4% in manufacturing and 5% in the ICT sector.

Table 6 shows the mean conditional inefficiencies by industry and the type of option schemes. As can be seen, in the ICT sector mean conditional inefficiencies vary in the range of 0.44-0.47 indicating that there is no clearly observable mean inefficiency difference between option and non-option firms. However, in the manufacturing sector our findings suggest that broad-based firms (0.28; 70 observations) may have higher mean inefficiency than selective (0.22; 236 observations) and non-option (0.21; 265 observations) firms, though the number of observations differs by the type of option scheme.

To examine whether option programs affect the mean and the variance of the inefficiency term, we use the estimator proposed by Wang (2002). Table 7 presents stochastic production function estimation results (with standard errors adjusted for intragroup correlation), when the mean and the variance of the inefficiency term are modelled as a function of option plans and firm size. Contrary to Table 4, where our option program indicators measure the presence of a plan, now the variable is *potential dilution*, measuring the potential size of an effective scheme in firm i in year t.[29] The following key findings emerge from Table 7. First, stochastic production frontier parameter estimates are in line with those presented in Tables 4. In addition, Wald tests clearly indicate constant returns to scale in production. Second, in both sectors the assumption that all parameters (constant excluded) are jointly zero is rejected for the variance of the inefficiency term, but not for the mean. However, the Wald test for the hypothesis that selective (*diluss*) and broad-based (*dilubb*) option scheme parameters are jointly zero both in the mean and in the variance of the inefficiency term is rejected in both sectors.[30] In sum, the tests support the parameterization of the mean and the variance of the inefficiency term. While informative, Table 7 does not provide an estimate of the magnitude of the effects of selective (*diluss*) and broad-based (*dilubb*) schemes on the mean and the variance of the inefficiency term u_{it}.

To provide a quantitative assessment of the marginal effects, Table 8 reports the marginal effects of the variables on $E(u_{it})$ and $Var(u_{it})$. The standard errors are bootstrapped results of 1,000 replications and significance levels are based on bias-corrected and accelerated intervals. The overall results support the view that especially broad-based schemes may affect the mean and the variance of the inefficiency term u_{it}. The results on $E(u_{it})$ show that an increase in the potential dilution of broad-based schemes is likely to increase production inefficiency. The average marginal effect is estimated to be 0.62, i.e. a one percentage point increase in the potential dilution of broad-based schemes increases firm technical inefficiency by 0.62%. Since

$\partial E(\ln(va)) / \partial dilubb = -\partial E(u) / \partial dilubb,$ the marginal effect on productivity would be about -0.62%. The average marginal effect of the potential dilution of broad-based schemes on $Var(u_{it})$ is positive, implying an increase in production uncertainty. Together these results would suggest, other things equal, that as time goes by broad-based scheme firms in the manufacturing sector may achieve lower and more uncertain productivity growth. For selective option schemes, we find no evidence that they affect firm inefficiency. Finally, our findings do not provide any empirical support for the view that stock option schemes reduce firm technical inefficiency in the manufacturing or ICT sector.

5. Conclusions

In this paper we study whether (i) firm-level technical inefficiency is higher in non-option than in option firms and (ii) whether the impact of options on firm technical inefficiency is related to the type of plan, i.e. whether the plan is broad-based or selective. We estimated stochastic production frontier models using novel panel data of Finnish publicly listed firms in the manufacturing and ICT sectors over the period 1992-2002. Our data enabled a careful investigation of the inefficiency effects of different types of option plans.

The key findings can be summarized as follows. First, the mean inefficiency estimates in the ICT sector are clearly higher than in the manufacturing sector. Efficient and inefficient firms exist in both sectors, but on average, mean inefficiency is higher in the ICT sector than in the manufacturing sector.

Second, our findings suggest that broad-based stock option firms in the manufacturing sector may have higher mean inefficiency than selective and non-option firms. On the contrary, in the ICT sector the mean inefficiency estimates do not indicate any difference between option and non-option firms.

Third, the quantitative assessment of the marginal effects supports the view that especially broad-based schemes in the manufacturing sector may affect the mean and the variance of the inefficiency term u_{it}. The results on the mean of the inefficiency $E(u_{it})$ show that an increase in the potential dilution of broad-based schemes increases production inefficiency in the manufacturing sector. Respectively, the average marginal effect of the potential dilution of broad-based schemes on the variance of the inefficiency term $Var(u_{it})$ implies production uncertainty in the manufacturing sector. These findings suggest that, other things equal, broad-based scheme firms in the manufacturing sector might achieve lower and more uncertain productivity growth as time goes by. Finally, we find no evidence that selective schemes affect firm inefficiency or the mean and the variance of the inefficiency term u_{it}. In summary, our findings do not provide any empirical support for the view that stock option schemes reduce firm technical inefficiency in the manufacturing or ICT sector.

Tables

Table 1. Development of stock option plans in Finland

Year	(1) No. of firms on the main list	(2) No. of firms in total	(3) No. of first option plan in this year	(4) No. of new option plans in this year	(5) No. of main list firms having option plans	(6) No. of firms having option plan	(7) HEX portfolio index, yearly changes[1]
1987	52	-	1	1	1 (1.9%)	1	-
1988	70	-	2	2	3 (4.3%)	3	-
1989	82	-	4	6	6 (8.5%)	7	-
1990	77	-	2	3	7 (9.1%)	8	-0.380
1991	66	-	3	4	9 (13.6%)	10	-0.113
1992	65	-	1	1	8 (12.3%)	11	0.077
1993	60	-	4	6	12 (20.0%)	15	0.657
1994	68	-	20	21	27 (39.7%)	34	0.164
1995	74	-	5	7	34 (45.9%)	38	-0.062
1996	73	-	3	9	34 (46.6%)	36	0.322
1997	82	115	12	22	40 (48.8%)	46 (40.0%)	0.273
1998	92	119	24	47	60 (65.2%)	69 (58.0%)	0.138
1999	102	137	21	41	77 (75.5%)	91 (66.4%)	0.541
2000	107	150	20	59	88 (82.2%)	113 (75.3%)	-0.242
2001	103	145	4	32	87 (84.5%)	112 (77.2%)	-0.191
2002	99	137	1	29	82 (82.8%)	101 (73.7%)	-0.150
Total			127	290			

[1] The portfolio index in trade-weighted average share returns, where a maximum weight assigned to one company is 10%. For years 1990-1995 we have used the general index, since the portfolio index is calculated only since 1996. Changes are in logarithmic scale.

Table 2. Summary statistics for the manufacturing sector

Variable	Name	Firm-year obs	Mean	Std. Dev.	Min	Max
ln(va)	Natural logarithm of value-added	571	18.27	1.79	14.95	22.37
ln(l)	Natural logarithm of employees	571	7.38	1.67	4.32	10.74
ln(k)	Natural logarithm of fixed capital	571	18.37	2.16	13.94	23.58
dilu*	Potential dilution in the range of (0,1); a proxy of option program size	298	0.0482	0.0462	0.0031	0.3069
diluss*	Potential dilution for selective stock option programs	228	0.0322	0.0227	0.0031	0.1109
dilubb*	Potential dilution for broad-based stock option programs	70	0.1003	0.0624	0.0369	0.3069
opt	Option program dummy	571	0.522	0.500	0	1
ssopt	Selective option program dummy	571	0.399	0.490	0	1
bbsopt	Broad-based option program dummy	571	0.123	0.328	0	1

All value measures are deflated using an industry-specific gross output deflator at 2000 constant Euros obtained from Statistics Finland. * Summary statistics for dilu, diluss and dilubb variables are only for those firms that have a stock option program. The data contains 571 firm-year observations regarding 62 firms.

Table 3. Summary statistics for the ICT sector

Variable	Name	Firm-year obs	Mean	Std. Dev.	Min	Max
ln(va)	Natural logarithm of value-added	243	17.13	1.50	13.56	22.16
ln(l)	Natural logarithm of employees	243	6.35	1.38	3.95	10.62
ln(k)	Natural logarithm of fixed capital	243	16.62	1.78	13.87	21.17
dilu*	Potential dilution in the range of (0,1); a proxy of option program size	139	0.0703	0.0452	0.0018	0.2138
diluss*	Potential dilution for selective stock option programs	64	0.0580	0.0377	0.0018	0.1872
dilubb*	Potential dilution for broad-based stock option programs	75	0.0807	0.0486	0.0184	0.2138
opt	Option program dummy	243	0.572	0.496	0	1
ssopt	Selective option program dummy	243	0.263	0.441	0	1
bbsopt	Broad-based option program dummy	243	0.309	0.463	0	1

All value measures are deflated using an industry-specific gross output deflator at 2000 constant Euros obtained from Statistics Finland. * Summary statistics for dilu, diluss and dilubb variables are only for those firms that have a stock option program. The data contains 243 firm-year observations regarding 32 firms.

Table 4. Stochastic production frontier estimates

	(1)	(2)	(3)	(4)
	Manufacturing Pooled MLE		ICT Pooled MLE	
constant	7.705 *** (0.156)	7.707 *** (0.157)	9.502 *** (0.408)	9.496 *** (0.410)
ln (labour)	0.709 *** (0.019)	0.710 *** (0.019)	0.853 *** (0.060)	0.856 *** (0.061)
ln (capital)	0.293 *** (0.012)	0.293 *** (0.011)	0.161 *** (0.042)	0.160 *** (0.043)
year	0.026 *** (0.006)	0.025 *** (0.006)	0.007 (0.010)	0.007 (0.010)
Parameters in the variance of v				
constant	-2.185 *** (0.519)	-2.238 *** (0.505)	-2.687 (1.920)	-2.699 (2.095)
ln (labour)	-0.182 ** (0.074)	-0.174 *** (0.072)	-0.167 (0.324)	-0.170 (0.354)
Parameters in the variance of u				
constant	-3.466 *** (0.997)	-3.889 *** (1.104)	-0.590 (0.654)	-0.964 (0.843)
ln (labour)	0.125 (0.116)	0.197 (0.131)	-0.018 (0.108)	0.051 (0.144)
opt	0.344 * (0.183)	-	0.622 ** (0.174)	-
ssopt	-	0.039 (0.193)	-	0.397 (0.425)
bbsopt	-	0.791 ** (0.326)	-	0.773 *** (0.169)
year	-0.037 (0.045)	-0.052 (0.048)	-0.064 (0.064)	-0.073 (0.063)
σ_v	0.175	0.176	0.155	0.156
σ_u	0.279	0.276	0.677	0.677
σ	0.330	0.327	0.695	0.695
$\lambda=\sigma_u/\sigma_v$	1.59	1.57	4.37	4.34
Log likelihood function	6.585	10.211	-124.114	-123.307
Finite sample corrected AIC	7.223	2.050	269.176	269.757
Wald test for constant returns to scale (p-value) Ho: $\beta_k+\beta_l=1$	0.88	0.85	0.96	0.69

The dependent variable is ln(value-added). Standard errors in parenthesis. ***, **, * statistically significant at 1%, 5% and 10% levels, respectively. We have 62 firms/571 observations in the manufacturing sector and 32 firms/243 observations in the ICT sector. SSOPT is a dummy variable for selective and BBSOPT is a dummy for broad-based option schemes, respectively. As a control group we use non-option firms.

Table 5. Conditional inefficiencies

Estimated inefficiencies $\hat{u}_{it}$	(1) Manufacturing Pooled MLE	(2)	(3) ICT Pooled MLE	(4)
Mean	0.217	0.213	0.451	0.449
Standard deviation	0.125	0.126	0.258	0.257
Minimum	0.040	0.041	0.050	0.051
Maximum	0.762	0.780	0.959	0.954

Conditional inefficiencies are based on the models presented in Table 4.

Table 6. Mean conditional inefficiencies by industry and the type of stock option scheme

	(1) Selective scheme firms	(2) Broad-based scheme firms	(3) Non-option firms	(4) Total
Manufacturing	0.218 (236 obs)	0.280 (70 obs)	0.190 (265 obs)	0.213 (571 obs)
ICT	0.445 (64 obs)	0.467 (75 obs)	0.438 (104 obs)	0.449 (243 obs)
Total	0.266 (300 obs)	0.377 (145 obs)	0.260 (369 obs)	0.283 (814 obs)

Conditional inefficiencies are based on the models presented in Table 4.

Table 7. Pooled stochastic production frontier ML estimates when parameterizing the variance and the mean of u

	Manufacturing	ICT
constant	7.791 (0.042) ***	9.570 (0.440) ***
ln (labour)	0.707 (0.042) ***	0.863 (0.052) ***
ln (capital)	0.286 (0.034) ***	0.150 (0.043) ***
year	0.023 (0.007) ***	0.003 (0.014)
Parameters in the mean of u		
constant	-17.634 (110.227)	-1.541 (2.786)
year	1.302 (8.040)	-0.423 (0.463)
diluss	-12.697 (68.423)	26.733 (25.529)
dilubb	-3.324 (33.001)	-10.269 (9.962)
ln (labour)	0.194 (1.006)	0.249 (0.288)
Parameters in the variance of u		
constant	0.873 (6.222)	0.573 (1.654)
year	-0.221 (0.062) ***	0.124 (0.077)
diluss	0.573 (11.964)	-5.838 (4.787)
dilubb	5.055 (1.722) ***	3.829 (4.233)
ln (labour)	0.065 (0.091)	-0.180 (0.190)
Parameters in the variance of v		
constant	-3.280 (0.194) ***	-3.494 (0.339) ***
Wald test for constant returns to scale (p-value) Ho: $\beta_k+\beta_l=1$	0.64	0.45
Wald test for joint significance of variables in the mean of u (without constant, p-value) Ho: *year, diluss, dilubb, ln(labour)* =0	0.99	0.45
Wald test for joint significance of variables in the variance of u (without constant, p-value) Ho: *year, diluss, dilubb, ln(labour)* =0	0.00 ***	0.00 ***
Wald test for joint significance of *diluss* and *dilubb* variables in the mean and variance of u (without constant, p-value) Ho: *diluss, dilubb* =0	0.03 **	0.046 **
Log pseudolikelihood	13.595	-112.599

The dependent variable is ln(value-added). Diluss is a selective and dilubb is a broad-based option scheme proxy variable, respectively. Standard errors in parentheses in the stochastic production frontier are robust (adjusted for intragroup correlation). ***, **, * statistically significant at 1%, 5% and 10% levels, respectively. We have 62 firms/571 observations in the manufacturing sector and 32 firms/243 observations in the ICT sector.

Table 8. Marginal effects on inefficiency

	Manufacturing	ICT
Marginal effects on $E(u_{it})$		
year	-0.0025 (0.285)	-0.0124 (0.0177)
diluss	-0.1907 (0.7153)	1.4733 (1.7382)
dilubb	0.6155 (0.2976) **	-0.0262 (1.4776)
ln(labour)	0.0129 (0.0102)	-0.0291 (0.0375)
Marginal effects on $Var(u_{it})$		
year	-0.0012 (0.0021)	0.0003 (0.0137)
diluss	-0.0499 (0.1941)	0.4539 (1.671)
dilubb	0.1780 (0.0893) **	0.1966 (0.7991)
ln(labour)	0.0036 (0.0028)	-0.0252 (0.0273)

Table reports sample means of marginal effects. Standard errors of marginal effects are bootstrapped results of 1,000 replications, statistical significant levels are based on bias-corrected and accelerated confidence intervals. ***, **, * statistically significant at 1%, 5% and 10% levels, respectively.

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Endnotes

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[2] Mäkinen (2001) describes the evolution of stock option programs in Finland. Jones, Kalmi and Mäkinen (2006a) study the determinants of option schemes adoption in Finland. They also summarise in more detail the evolution of options and discuss the institutional background in Finland.

[3] See also Wang (2002) and Bottasso and Sembenelli (2004).

[4] Although the Finnish economy was only the 47th largest in the world in 2003, it is an extremely interesting case. First, The World Economic Forum (WEF) found Finland the most competitive in the survey of 104 economies in 2003-2004. More surprisingly, Finland has subsequently held the top position in three out of the last four years. Second, Transparency International ranks Finland as the world's least-corrupt country for the fifth consecutive year. Third, during the period 1992-2002 the great majority of publicly listed Finnish firms adopted their stock option plans. This enables us to study the effects of options in a place where their use was previously rare, such as in Finland.

[5] For a more detailed description see e.g. Jones, Kalmi and Mäkinen (2006a).

[6] For more detailed discussion about the Great Finnish Depression during 1990-1993 see e.g. Kiander and Vartia (1996), and Honkapohja and Koskela (1999).

[7] Firms may adopt schemes for different reasons. For example, the shareholders of a firm may prefer to broaden schemes to a larger set of employees, or there is a need to change the terms of a scheme for some reason.

[8] This is done to increase the number of observations in the data.

[9] We are aware that few unlisted Finnish firms have also adopted option schemes, at least during the bull market in the end of 1990s. Unfortunately, no information on these firms and option schemes was unavailable. We can only roughly conclude that it is perhaps more probable to find these programs within the ICT sector than within other sectors. We believe, however, that the number of these unlisted stock option firms is small, since option schemes works properly only in situations where the value of shares can be assessed in the stock market. Also, in order to study the impacts of stock option programs with public data, our data seem to be a reasonable choice.

[10] Our classification is different from Kroumova et al. (2006), who use a 50 % threshold as a criterion for broad-based schemes. Our data do not include this information, but they have the important advantage of being derived from publicly reported sources that must be externally verifiable, rather than from confidential surveys.

[11] We also interviewed Mr. Erkki Helaniemi, a partner at Alexander Corporate Finance, an investment bank, who has been personally involved in setting up dozens of option schemes. He confirmed that there are dramatic differences in the participation rates for option schemes, depending on eligibility.

[12] Unfortunately, we do not have information on stock option program details, such as exercise prices to calculate Black-Scholes values.

[13] Excellent literature surveys are Bauer (1990), Greene (1993; 1997), and Kumbhakar and Lovell (2000).

[14] For example, the linear fixed effects estimator captures all fixed effects between firms potentially making firm-specific inefficiency indecomposable from heterogeneity among firms. Moreover, at least one producer is assumed to be 100% technically efficient, and other firms' inefficiency is measured relative to this fully efficient producer. The random effects estimator suffers from the assumption that firm-specific inefficiency is the same in every year. For short panels this may be an appropriate assumption, but in longer panels this is likely to be problematic. Other drawbacks of the random effects estimator are that heterogeneity between firms is absorbed into the inefficiency term, and it is assumed that the inefficiency term is uncorrelated with other explanatory variables. Thus, as argued by Greene (2005), both traditional linear panel estimators previously used in the stochastic frontier literature may be seriously distorted due to blending of inefficiency and heterogeneity in the same term.

[15] The noise component v_{it} captures measurement errors and production function misspecification effects, whereas u_{it} is related to technical inefficiency.

[16] It is also possible to obtain confidence intervals for the point estimates of technical inefficiency, but we do not examine this issue here. For more details see Horrace and Schmidt (1996), and Bera and Sharma (1999).

[17] See Kumbhakar and Lovell (2000) for detailed discussion on how to account for exogenous influences in the one- and two-step approaches. Also, according to the Monte Carlo studies conducted by Schmidt and Wang (2002), the one-step modelling approach is more favourable than the two-step approach, where inefficiencies and exogenous effects are estimated sequentially.

[18] As a novel contribution to the stochastic frontier literature, Greene (2002a, 2005) greatly extends a simultaneous accounting of heteroskedasticity and inefficiency by proposing e.g. a new "true" fixed effects framework that more explicitly follows stochastic frontier modelling foundations applied frequently in cross-section frontier models. See Prentice and Gloeckler (1978), Sueyoshi (1993), and Greene (2001, 2002a) for a formal derivation of the estimators. See also LIMDEP's manual and Greene (2002b, 2005).

[19] See Wang (2002).

[20] On theoretical grounds firm value added is a preferable measure to sales (i.e. as a proxy for firm output), since value added does not include intermediate inputs that are purchased from other firms.

[21] The following issues have influenced our empirical strategy. First, we assume a Cobb-Douglas form of technology, since it has been used frequently in the related productivity literature, such as the evaluation of the effects of ESOPs on firm productivity (e.g. Jones and Kato, 1995) and analysing the effects of stock options on firm productivity (e.g. Conyon and Freeman, 2004; Jones, Kalmi and Mäkinen, 2006b). Second, although the Cobb-Douglas functional form is more restrictive than other functional forms, such as the translog, we prefer the Cobb-Douglas production function, since the number of the translog production frontier model estimations worked poorly, i.e. the maximum likelihood estimators did not converge in the estimations. Third, since we do not have information on the detailed terms of option schemes, such as exercise prices, we must bypass some potentially important issue. For example, presumably the terms of option schemes differ among firms, which may affect performance effects. Thus, when an option scheme is substantially out of the money (i.e. a current stock price is substantially below an exercise price), options may not provide strong incentives for employees and managers to improve their performance.

[22] A priori we conducted several fixed effects estimations by LIMDEP where we modelled the mean and the variance of the inefficiency term. Unfortunately, all models behaved extremely poorly, even when we tried the stratification method. As noted in LIMDEP's manual pp. E24-27, fixed effects formulations, especially based on the Newton's method, can be "extremely problematic in all but the most favourable of cases". Therefore we can only report here results based on the pooled ML model.

[23] According to Bottasso and Sembenelli (2004), unaccounted heteroscedasticity in the u_{it} leads to biased estimates of the production frontier parameters and technical efficiency, whereas unaccounted heteroscedasticity in the v_{it} leads to biased estimates of technical efficiency.

[24] The mean measures the expected value of technical inefficiency, whereas the variance measures production uncertainty (Bera and Sharma, 1999).

[25] The estimation approach has a lot in common to Bottasso and Sembenelli (2004), who provide evidence on the relation between identity of ultimate owners and technical inefficiency by estimating stochastic production frontiers on Italian manufacturing firms.

[26] Note that we use here a different estimator. The reason is that the model, kindly provided by Hung-Jen Wang, utilises STATA's maximum likelihood routines and assumes that same the z affects both the mean and the variance of u_{it}.

[27] Wang (2002) underlines that the marginal effects on the conditional mean and variance of u_{it} are almost intractable, particularly when the variances of u_{it} and v_{it} are modelled.

[28] Based on a result from Barrow and Cohen (1954), $m_1^2-m_2>0$.

[29] The reason is that we had major problems in convergence of the ML estimator when using option program dummy indicators.

[30] We also specified models (not reported here) where the variance of v_{it} was modelled as a function of firm size. All estimated models and performed Wald tests indicated that v_{it} is not heteroskedastic with respect to firm size.

CEO Compensation, Firm Size and Firm Performance: Evidence from Finnish Panel Data[1]

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Abstract:
This paper examines how CEO pay is related to firm size and to firm performance in Finland by using new individual-level compensation data in 1996-2002. We find robust evidence that CEO average compensation has increased substantially between 1996 and 2002. For example, the ratio between CEO and industrial worker mean total compensation was 7 in 1996, peaked at 24 in 2000, and thereafter dropped to 13 in 2002. We argue that the change in CEO compensation, and especially in total compensation, is highly related to changes in stock market measures of firm performance. Our shareholder wealth measure suggests that the salary and bonus change in CEO wealth is €6.84 per €1,000 change in shareholder wealth. Respectively, the total compensation change is €21.85 per €1,000 change in shareholder wealth. We find no evidence on the contemporaneous link between a change in CEO compensation and change in ROA% (Return on Assets). However, one-year lagged accounting and stock market based firm performance measures are associated with the change in CEO total compensation. In line with previous studies, our findings suggest that pay-for-firm size elasticity is close to 0.3. We also find interesting corporate governance findings. First, the share of foreign ownership is positively and statistically significantly associated with the level of compensation. Also, foreign ownership parameter estimates are about three times larger for total compensation than for salary and bonuses in most specifications. Second, ownership concentration, as measured by the voting share of a largest shareholder, is negatively related to the level of compensation, but only in the pooled model. Third, the size of the board is positively related to the level of compensation, especially to the level of base salary and bonus.

Keywords: CEO compensation, pay for performance, stock options, firm size

JEL-codes: J33, M52, L25

1. Introduction

After Enron Corp's financial frauds and bankruptcy in December 2001, chief executive officer (CEO) remuneration has been a popular topic of public debate in a number of countries (e.g. The Economist, 1999; 2000; Krugman, 2002; and Samuelson, 2003). Enron Corp's scandal was not a unique incidence, and afterwards we have witnessed, for example, the accounting and compensation scandals of Arthur Andersen, AOL Time Warner, Dynegy, Merck, Qwest, Tyco, WorldCom and Xerox in the U.S., Parmalat in Italy, Dutch-based Royal Ahold, and Swedish Scandia. Shareholders of these firms have lost billions of dollars, and the scandals are said to be related to an increased use of equity-based incentive components in CEO compensation, such as stock options.

With the increased use of stock options in CEO compensation packages, CEO average compensation has increased substantially, especially in the U.S. For example, the CEO average pay at the largest companies in the U.S. was 40 times that of the average worker a generation ago, but in 1999 it was 475 as much. In 2002, the pay of top American CEOs was still over 400 times average earnings, but in 2004 the figure is estimated to have fallen close to 160 (The Economist, 2005). On the contrary, the estimated ratios are considerably smaller in Europe, ranking from 11 in the Switzerland to 24 in the U.K. times that of average employees in 1999 (The Economist, 2000).

The differing interests between shareholders and top executives is not a recent notion in economics. In fact, already Adam Smith (1776) suggests: ... "*What are the common wages of labour depends everywhere upon the contract usually made between those two parties, whose interests are not the same. The workmen desire to get as much, the masters to give as little as possible* " ... "*The directors of [joint stock] companies,* *however, being the managers of other people's money than of their own, it cannot be expected that they should watch over it with the same anxious vigilance [as owners]... Negligence and profusion, therefore, must prevail, more or less, in the management of the affairs of such a company".* Later Berle and Means (1932) propose that the separation of ownership and control in a modern corporation may introduce a principal-agent problem due to asymmetric information between shareholders and executives.[2] Although it is unjustified to categorise executives' behaviour as a group, asymmetric information may encourage opportunistic and ineffective behaviour, which in turn may lead to decreases in shareholder value.[3] Thus, executives' compensation packages may be revised so that managers have monetary incentives to exert more effort and take actions that are mutually beneficial to both them and shareholders.

This study examines CEO compensation in Finland over the period 1996-2002. By providing new evidence from a very different institutional context than the U.S. and the U.K., we hope to increase our understanding of CEO compensation practices across different countries. In particular, we follow previous empirical studies in the literature by exploring CEO pay-for-firm size elasticity and CEO pay-for-firm performance sensitivity. We estimate several empirical specifications where we control for industry of the firm, CEO age, the size of the board, the voting share of the largest shareholder and the share of foreign ownership, since all these variables may affect the level and the changes of CEO compensation.

Unlike in the U.S., where publicly listed firms are required to disclose detailed information on the top five executives' compensation, Finnish listed firms typically disclosed only aggregated total compensation of the CEO and the board of directors in 1996-2002. To avoid measurement error biases associated with aggregated data, we utilise new, hand-collected individual-level CEO compensation data obtained from Finnish tax authorities' registers. In this annual compensation data pay is divided into different categories allowing us to separate CEO compensation from different origins. Our data also contain a variable that includes executives' revenues from exercised stock options during a given year.

The key finding is that CEO average compensation has increased substantially between 1996 and 2002. For example, the ratio between CEO and industrial worker average total compensation was 7 in 1996, peaked at 24 in 2000, and thereafter dropped to 13 in 2002. In real terms the mean salary and bonus of CEOs was €166,000 (median €147,000) in 1996, whereas it was €280,000 in 2002 (median €208,000). The percentage increase from 1996 to 2002 is 69% (median 41%). Respectively, CEO mean total compensation increased from €180,000 to €357,000 (98%), whereas median total compensation increased from €155,000 to €233,000 (50%). This development is related to the stock market boom and bust in the Helsinki Stock Exchange, since changes in CEO compensation, and especially in total compensation, are highly related to changes in stock market-based measures of firm performance. Our shareholder wealth measure, close to that of Jensen and Murphy (1990), suggests that the salary and bonus change in CEO wealth is €6.84 per €1,000 change in shareholder wealth. Respectively, the total compensation change is €21.85 per €1,000 change in shareholder wealth, being likely upward biased due to a few large stock option exercises. Moreover, the estimated "semi-elasticity" of CEO salary and bonus with respect to stock returns is 0.09, and 0.28 for total compensation. We did not find statistical evidence on a contemporaneous association between the change in CEO compensation and change in ROA% (Return on Assets), an accounting-based measure of firm performance. However, changes in one-year lagged performance measures, both accounting and stock market-based, can be associated with the change in CEO total compensation.

Another key finding is that the pay level of CEOs is related to firm size: pay-for-firm size elasticity does not deviate substantially from 0.3, after controlling for CEO age, industry of the firm, ROA% and three corporate governance indicators, namely the size of the board, the voting share of the largest shareholder, and the share of foreign ownership.[4] Also interestingly, from a corporate governance perspective, the CEO pay level is positively related to foreign ownership and the size of the board, but negatively to a dominant shareholder's ownership.

The paper is organised as follows. Section 2 briefly describes the principles of corporate governance mechanisms in a publicly listed firm. It also summaries major institutional and corporate governance changes in Finland in the 1990s. Section 3 summarises the relevant previous literature. In section 4 we describe the data and present the empirical models. Section 5 presents the estimation results. Finally, section 6 concludes.

2. Corporate Governance Principles and Institutional Changes

2.1 Corporate Governance Principles

In a publicly listed firm the principles of corporate governance mechanisms can be factorised into internal and external governance.[5] From shareholders perspective, internal corporate governance can be associated with shareholders' annual general meeting and the active governance done by shareholders' representatives, the board of directors. External corporate governance is strictly defined as an increasing threat of takeover (e.g. when a firm's performance is inferior), but more broadly understood it may contain the actions of all outside stakeholders, competitors and regulators. Internal governance mechanisms may consider as primary factors to detect and to prevent corporate scandals, and if these mechanisms fail, then external governance is likely to intervene in corporate control, at least in well-functioning stock markets.[6]

As a solution to mitigate the principal-agent problem, shareholders' may directly monitor executives' behaviour. However, if ownership is highly dispersed, it is ineffective for all shareholders separately to carry out monitoring and governance tasks. Besides ineffectiveness, a highly dispersed ownership exposes shareholders' individual monitoring to free-rider problems. Obviously, shareholders can do much better, if select the board of directors to represent and to ensure their financial interests. Since a modern firm has a full legal capacity, the principal tasks of the board of the directors are illustrated in corporate law and regulations, and nowadays more often in corporate governance recommendations.[7]

As a second potential solution to mitigate the principal-agent problem, shareholders may utilise incentive-based remuneration, such as accounting-based bonuses, restricted stocks and stock options, in order to link financial interests between shareholders and executives. The key difference between accounting-based bonuses and equity-based instruments is that the value of a firm's stock is determined outside of executives' direct control if a stock market is well-functioning and efficient. On the contrary, a badly implemented accounting-based bonus plan may be adjusted by executives, at least to some extent. Also, accounting-based bonuses focus on a firm's annual and past performance, whereas stocks and stock options reflect more a firm's future growth potential. With a carefully implemented and adopted equity-based incentive programs, executives' attention should be on selecting and implementing actions that increase the firm's total value in the long-run.[8] We also underline that corporate governance practices and executive compensation policy principles are closely related issues. Without coherent corporate governance practices within a firm, executives' compensation packages may be implemented badly, potentially spurring a decrease in the shareholders' wealth, e.g. via corporate scandals.[9]

2.2 Institutional Changes

During the 1990s Finland experienced substantial institutional changes in corporate governance, foreign ownership, industrial relations and globalisation of firms. The Finnish *corporate governance* system in listed firms was very much bank-centred and resembled the German stakeholder system at the end of 1980s. During this time moderate accounting-based bonuses, if any, were the only performance pay component in CEO compensation. Financial institutions owned around 25% of the value of shares in the Helsinki Stock Exchange. Bank loans were the most significant source of external funding for listed companies.[10] On the other hand, at the end of the 1980s, the stock market was booming and the number of listed firms was at a record high. This, however, ended in the early 1990s, when Finland suffered the most severe depression in any OECD country since the Second World War. For example, during 1990-1993 unemployment soared close to 20% and GDP plummeted by 14%.[11] This caused e.g. a significant change in the financial markets: the value of bank loans dropped significantly, as did share prices on the Helsinki Stock Exchange. After the devaluation of the Finnish currency in 1991 and its floating in 1992, the stock market and the economy started to recover, but bank lending continued to decline throughout the 1990s. Turnover on the Helsinki Stock Exchange grew dramatically during the 1990s (although this is partly because of the growth of Nokia) and the number of firms listed also increased significantly, especially in the late 1990s and 2000. Nowadays stock markets are much deeper, more informative and more transparent. At the same time, both monitoring of insider trading and legal punishments have become stricter. During the last 10-15 years Finland has shifted from a system of bank-based financial intermediation closer to a market-based Anglo-American system. As a part of this institutional change publicly listed Finnish firms have extensively adopted stock option schemes in the 1990s.[12]

Another important institutional change is the increase of *foreign ownership* in publicly listed Finnish firms. The Finnish stock market was fully opened to foreign investors only in 1992, but today foreigners are the largest ownership group (although again this is largely because of Nokia). By 2000, foreign ownership had increased to 53%, while ownership by domestic financial institutions had dropped at the same time from 20% to 4% (Hyytinen, Kuosa and Takalo, 2003). According to Barca and Becht (2001), an increase of foreign ownership has triggered changes in corporate governance in many European countries. This is the case also in Finland, where, according to Tainio and Lilja (2003), increase in foreign ownership has contributed to the transformation of Finnish business towards a more competitive and open culture, where shareholder value is given a high priority. Moreover, foreign owners may have also played a major role when the use of stock options increased during the 1990s.

Third, turning to *industrial relations*, we observe both continuity and change. Traditionally, Finland has been seen as a highly egalitarian society, which from an industrial relations perspective is characterised by high labour taxes, an extensive public sector, and small wage dispersion. Consensual collective bargaining and centralised income agreements have continued as the norm for decades. Since the late 1960s, the unionisation rate of the workforce has been around 70-80%, and collective agreements are typically binding also for non-union workers or workplaces. Wage increases consist of a collectively agreed element that is typically economy-wide. In addition, firms can adapt their internal wage structures according to their financial possibilities. Throughout the 1990s, profit-sharing and other forms of performance-related pay have become common compensation methods throughout the economy (Kauhanen and Piekkola, 2002). Forms of performance-related pay are not negotiated in collective bargaining rounds, but employers can decide on their use unilaterally. The widespread use of performance-related pay, as well as the popularity of stock options, represents a change in industrial relations. However, despite the increase of performance and equity-related compensation within companies, there has been only a moderate rise in wage dispersion in Finland[13], especially by international standards.

Finally, when focusing on the *globalisation of large Finnish firms* during the 1980-1990s, we observe dramatic change in the number of firms' employees abroad. For example, the share of ten major Finnish firms' employees abroad was about 15% of personnel in 1983, whereas in 2002 the share was over 60%. This huge increase reflects the fact that in the early 1980s Finnish firms mainly exported their products from Finland. However, gradually firms adopted more complex business practices abroad, such as their own production units. The value of inward foreign direct investments increased in the end of 1990s due to active cross-border mergers and international acquisitions of Finnish firms.[14] In sum, changes in globalisation of large Finnish firms indicate that their business environment is nowadays much larger and more complex than in the 1980s.

3. Previous Literature

The previous empirical literature on CEO compensation is multidisciplinary. Various academics from economics, finance, accounting and management fields have contributed to the current state of the literature. The vast majority of this extensive research has emerged during the last 25 years, since before the 1980s only a handful of CEO compensation studies were published (e.g. including works by Roberts, 1956; Baumol, 1959; Lewellen and Huntsman, 1970; and Becker, 1975). Since then research has been conducted *in economics* (e.g. Jensen and Murphy, 1990; Rosen, 1990; Gregg, Machin and Szymanski, 1993; Conyon and Leech, 1994; Main, Bruce and Buck, 1996; Vittaniemi, 1997; Hall and Liebmann, 1998; Murphy, 1999; Conyon and Murphy, 2000; Murphy, 2000; and Kato and Kubo, 2005), *in finance* (e.g. Yermack, 1997), and *in accounting and management* (e.g. Finkelstein and Boyd, 1998; Gomez-Mejia and Wiseman, 1997; and Randøy and Nielsen, 2002). We next survey the directly related studies.

Lewellen and Huntsman (1970) analyse data on 50 U.S. firms at three-year intervals from 1942 to 1963. They find strong evidence that top executives' compensation is heavily dependent upon the generation of firm profits. Their results also indicate that firm accounting-based profits and stock market values are substantially more important in the determination of executive compensation than are firm sales.

Jensen and Murphy (1990) use CEO compensation data on a sample of 1,295 U.S. firms from 1974 to 1986. They estimate pay-for-performance models in first-differences to study how change in CEO compensation is related to change in shareholders' wealth. They find that CEO pay-for-performance sensitivity has been modest and it has fallen in real terms from the 1930s: *"... on average, corporate America pays its most important leaders like bureaucrats. ... The total change in all CEO wealth is $3.25 per $1,000 change in shareholder wealth for the full sample, $1.85 for large firms, and $8.05 for small firms. The largest CEO performance incentives come from ownership of their firm's stock."*

Rosen (1990) surveys several independent empirical studies on CEO pay-for-firm performance. Based on the evidence from these studies, he concludes that the effect of stock returns on log compensation is in the range of 0.10-0.15.

Gregg, Machin and Szymanski (1993) focus on the relationship between the wage of the highest paid director and firm performance with a U.K. data sample of 288 large listed firms from 1983 to 1991. They find evidence that, in terms of share returns over the whole period, the relationship between the top director's pay and firm performance is very weak in the period. However, after splitting the data into two sub-periods, i.e. 1983-1988 and 1989-1991 (recession period), they find a positive but small pay-for-performance relationship for the first sub-period, but not for the second. When focusing on the link between the top director's pay and firm size, they argue that growth in the top director's pay is strongly correlated with the growth of firm size: a 50% increase in sales leads to a 10% increase in the top director's pay.

Conyon and Leech (1994) examine the determinants of the top director's salary and bonus with a sample of 294 large U.K. listed firms in 1983-86. They find a positive but very small pay elasticity with respect to firm performance. For the median director, a 10% increase in shareholder wealth corresponds to an increase in compensation of 375 pounds. Another key finding is that ownership control and concentration decrease the level of the top director's pay, but these variables do not affect the growth of pay.

Main, Bruce and Buck (1996) use U.K. panel data for 60 firms from 1981 to 1989. They find that because of stock options there is a statistically significant relationship between the wage of the highest paid executive and firm performance. For example, a 10% increase in shareholder wealth increases the top paid director's compensation about 9%. The sensitivity of top executive compensation with respect to firm performance is greater than in the previous U.K. studies, since they have also taken into account information on stock options.

Hall and Liebman (1998) use a 15-year panel of CEOs in the largest U.S. firms from 1980 to 1994. They argue that CEO compensation is highly responsive to firm performance if the value changes of CEO stock and option holdings are taken into account in the empirical analysis. For example, the median elasticity of CEO compensation with respect to firm market value is 3.9 for 1994, which is about 30 times larger than previous estimates that rely on salary and bonus changes alone. They also argue that CEO mean (median) compensation increased by 207% (146%) in real terms between 1980 and 1994. Perhaps more importantly, virtually all of this increase is attributable to changes in the value of CEO holdings of stock and stock options. When using an analogous measure to that of Jensen and Murphy (1990), in 1994 the total change in CEO wealth is $5.25 per $1,000 change in shareholder wealth. Although this degree of sensitivity may appear modest, Hall and Liebman show that CEO wealth may change millions of dollars for a typical change in firm value.[15] Thus, they conclude that CEO compensation is strongly related to the success of the companies they manage.

A majority of the previous empirical CEO compensation studies have been conducted either in the U.S. or in the U.K., mainly due to a better availability of data on CEO compensation.[16] There has also been interest in CEO compensation research in other countries recently. For example, Randøy and Nielsen (2002) examine the relationship between firm performance, corporate governance and CEO compensation in Sweden and Norway in 1998, by using data on 120 Norwegian and 104 Swedish publicly listed firms. The evidence based on cross-sectional estimates indicates a statistically significant and positive relationship between the size of the board and CEO compensation, foreign board membership and CEO compensation, and firm market capitalisation and CEO compensation. On the contrary, they do not find evidence that CEO compensation is statistically related to firm performance.

Kato and Kubo (2005) examine the link between CEO compensation and firm performance in Japan by using new panel data from 1986 to 1995. They find evidence that CEO cash compensation is sensitive to firm performance, especially for accounting-based measures of firm performance. However, stock market-based measures of firm performance seem to be a less important factor in CEO compensation. One reason may be the fact that until 1997 executives' stock options were banned in Japan, except at small venture capital companies.

In Finland, Vittaniemi (1997) has studied the relationship between CEO compensation and firm performance previously. He uses panel data on 48 listed and 70 non-listed firms in 1989-93 (5 years) and estimates separate models for listed and non-listed firms. As a firm performance measure he uses once lagged variables. He finds a significant pay-for-performance relationship in listed firms, but in non-listed firms the relationship is less important. Contrary to Vittaniemi, we focus only on listed firms, since we pay special attention to stock option compensation, which can work properly only in situations where the value of shares can be assessed in the stock market.[17] In addition, the time period in Vittaniemi's study is very exceptional in the Finland's economic history, since in 1990-93 Finland suffered the most severe depression in any OECD country since the Second World War.

When turning to the empirical research on the relationship between CEO compensation and firm size, an interesting and a well-documented finding is the relative uniformity of CEO pay-for-firm size elasticity estimates. For example, Baker, Jensen and Murphy (1988) report elasticities in the range of 0.25-0.35, when summarizing the U.S. Conference Board data on the link between CEO cash compensation and firm sales from 1973 to 1983. This finding is supported by Rosen (1990), who summarizes a variety of studies for different time periods in the U.S. and the U.K. He finds some variation in CEO pay-for-firm size elasticities, but "*...the relative uniformity of estimates across firms, industries, countries, and periods of time is notable and puzzling because the technology that sustains control and scale should vary across these disparate units of comparison. The estimated elasticities for all companies are not significantly different from 0.3.*" Recently Conyon and Murphy (2000) estimate CEO pay-for-firm size elasticities with the data on the U.K. and the U.S. firms in 1997. Their findings support "the near uniformity elasticity hypothesis $\beta=0.3$" for the U.S. ($\beta=0.3$), but not for the U.K. ($\beta=0.2$) firms.

Although the previous empirical studies commonly report an almost uniform 0.3 point estimate for CEO pay-for-firm size elasticity, the studies do not explain what might be a possible reason behind this phenomenon. To the best of our knowledge, we are unfamiliar with any theoretical study that might explain why the point elasticity estimate is near 0.3 across different firms, industries, times and countries. The only explanation that we are familiar with is Davidson Consultants' (1984) "Wage and Salary Administration in a Changing Economy", as noted in Baker, Jensen and Murphy (1988). It interestingly describes how a consulting firm sets a CEO pay-for-firm size elasticity coefficient: "*The general rule is that as sales volume doubles, executive pay increases by one-third.*" If Davidson Consultants' "general rule" presents a common practice among compensation consultants, it may explain surprising commonalities in elasticity point estimates across firms, industries, times and countries.

4. Data Description and Empirical Models

4.1 Data Description

We have combined several data sets for this study. First, CEO annual compensation data are hand-collected from the Finnish tax authorities' registers. The data is not a random sample, but we have used all the feasible information on CEOs to construct as large an individual-level compensation data set as possible.[18] Second, firm-level financial statement data are compiled from Balance Consulting, a Finnish consulting firm. Third, we use information on the firms' foreign ownership and market values, based on the data from the Helsinki Stock Exchange. Fourth, the largest shareholder's ownership and the size of the board data are hand-collected from the Pörssitieto-handbooks[19] and the firms' annual reports. Fifth, the data on stock returns were kindly provided by the department of finance and accounting from the Helsinki School of Economics (originally from the Helsinki Stock Exchange). Finally, all nominal monetary variables are deflated to real Euros of the year 2000 using gross-output based industry deflator, published by Statistics Finland.

The great benefit of our individual-level compensation data is that it disaggregates CEO's salary and bonus by origins.[20] The data contain information on CEO's taxable benefits, such as company car and other perquisites. The other perquisites item is especially interesting, since it includes the taxable income of exercised stock options in a given year for the years 1996-1999.[21] For more recent years from 2000 onwards, the data explicitly show the taxable income of exercised options.[22]

When assessing firm performance, we use both accounting and stock market-based measures, since both have been used in the previous literature. In economics and finance, most CEO pay-for-performance studies use stock market-based measures. In contrast, according to Joskow and Rose (1994), studies in the accounting literature typically use accounting-based or both measures.

As a stock market-based firm performance measure we use two variables that have been used previously. The first is annual stock return, which is based on a firm's continuously-compounded daily stock returns, i.e. $\ln\left[\left(p_t + d_t\right) / p_{t-1}\right]$, where p_t is the price of a firm's share in the last trade in period t, p_{t-1} is the last trading price in period $t-1$ and d_t is the dividend.[23] To make our stock market measure comparable to Jensen and Murphy (1990), our second stock market-based firm performance measure is shareholder wealth, i.e. $\ln\left[\left(\left(p_t + d_t\right) / p_{t-1}\right) * V_{t-1}\right]$, where V_{t-1} is firm market value in the beginning of a period.[24] As an accounting-based firm performance measure we use ROA% (Return on Assets), since this has been used previously.[25]

Table 1 summarizes CEO compensation data in real 2000 Euros over the period 1996-2002. The number of CEOs varies between 43 (in 1996) and 82 (in 2000).[26] Table 1 suggests that the level of CEO compensation has increased substantially between 1996 and 2002. For example, CEO mean salary and bonus in real terms was €165,878 (median €147,113) in 1996, whereas in 2002 it was €279,733 (median €207,856). In percentages, the increase was 69% (median 65%). Respectively, mean total compensation was €180,190 (median €155,142) in 1996, whereas it was €356,863 (median €232,750) in 2002. In percentages, the increase was 98% (median 50%) indicating a few large stock option exercises. When examining the compensation increases of the cohort 1996, the corresponding increases were 74% for both mean and median salary and bonus. Respectively, the percentage change in mean total compensation was 105% and 74% for median total compensation.

The growth trend in our CEO compensation data diverges clearly from the development of industrial workers' average compensation. For example, industrial worker mean total compensation in Finland was €24,793 in 1996, whereas in 2002 it was €27,660. In percentages, the increase was 12% between 1996 and 2002, which is a fairly moderate increase compared to the growth of CEO compensation. Moreover, the ratio of CEO mean total compensation and industrial worker mean total compensation was only 7 in 1996, but 24 in 2000, mainly due to executives' exercised options during the stock market boom. Thereafter the ratio has fallen to 13 in parallel with the stock market bust in 2000-2002. Our estimated ratio of 16 in 1999 exceeds somewhat that reported for Germany (13), Sweden (13) and France (15), but is little less than in the Netherlands (17), Spain (17) and Belgium (18).[27] These "European ratios" are substantially smaller than in the U.S. (475) in 1999.[28]

Table 1 also shows some other interesting issues. First, CEO compensation distributions are clearly skewed to the right, being a consequence of the fact that pay levels are higher in large firms. This is a well documented observation from the previous literature (e.g. Murphy, 1999; Conyon and Murphy, 2000).

Second, although CEO mean compensation increased substantially over the period 1996-2002, at the same time firms' mean EBIT (earnings before interests, taxes and extraordinary items) has increased even more in real terms (189%).[29] However, annual percentage changes in compensation and EBIT diverge in some years. In 1997, 1998 and 2000 both increased from the previous year, but in 1999 mean EBIT increased about 6%, whereas mean salary and bonus decreased -0.7%. In 2001, we see a drop of -42% in mean total compensation, when both the HEX portfolio index (-19%) and mean EBIT (-18%) sank. Surprisingly, however, mean salary and bonus increased already by

17% in 2001. In 2002, the development of CEO compensation and firm performance tends to be mixed: mean salary and bonus (9%) and mean EBIT increased (14%), but mean total compensation decreased (-5%).

Third, there seems to be a great variation in the yearly growth rates of mean and median salary and bonus, although this may be partially explained by the variation in the number of CEOs in a given year. Maen salary and bonus increased in 1998 (17%) and 2001 (17%), but decreased in 1999 (-0.7%). Respectively, median salary and bonus increased in 1997 (19%), 1998 (12%) and 2000 (10%), but decreased in 1999 (-9%).

Fourth, percentage changes in mean total compensation appear to be larger than changes in mean salary and bonus. For example, CEO mean total compensation increased 53% in 2000, then decreased 42% in 2001 the reason being the stock market slump that degraded the value of stock options. However, the number of CEOs that exercised stock options increased from 1 (1996) to 11 (2002) corresponding to a relative change from 2% (1996) to 16% (2002).[30] In parallel with the increase in the number of CEOs that exercised stock options, mean values of exercised options have exploded, as can be seen from the last row of Table 1. For example, in 1997 the mean value of exercised options was €369,137 (in real terms), in 2000 a record high €3,787,800, and even in the stock market bust years 2001-2002 over €900,000. At the same time the sum of all exercised options per year as a percentage of the sum of CEOs' total compensation per year has jumped from 2% (1996) to 43% (2002). This clearly indicates that a few CEOs have gained substantially from stock option-based compensation in Finland between 1996 and 2002.

[Table 1 about here]

Table 2 presents some summary statistics in 2000. CEO mean (median) salary and bonus is €220,549 (€193,495), whereas CEO mean (median) total compensation is €646,845 (€205,891). The firm size distribution, measured by sales, is skewed to the right (mean sales €1,370,000,000 is much larger than median sales €251,000,000). The median size of the board is 6 members, and CEO median age is 51 years. The mean of foreign ownership is 20.5%, which is a little less than the mean share of the dominant shareholder's ownership 22.1%.

[Table 2 about here]

Table 3 compares CEO mean total compensation in Finland, Sweden and Norway.[31] Table 3 suggests that, in 1998, mean total compensation is higher in Finland than in Norway or in Sweden, mainly because of large option gains of 8 Finnish CEOs. When we drop these CEOs from the comparison, mean total compensation in Finland (€185,000) is still higher than in Norway (€162,000), but lower than in Sweden (€280,000).

[Table 3 about here]

4.2 Empirical Models

The best-documented empirical finding is the relative consistency of the relationship between *CEO pay and firm size.* To follow the previous pay-for-firm size elasticity studies (e.g. Baker, Jensen and Murphy, 1988; Rosen, 1990; Murphy, 1999; Conyon and Murphy, 2000), we first estimate by OLS separate cross-section models in 1996-2002 for the following loglinear equation:

(1) $\ln\left(CEO\ salary\ and\ bonus_i\right) = \alpha + \beta \ln(Sales_i) + \varepsilon_i$, i=1,...,N, $\varepsilon_i \sim iid(0,\sigma^2)$.

In Equation (1) subscript *i* indexes individual CEOs at firm *i*. We estimate Equation (1) separately for CEO salary and bonus, and for CEO total compensation. In addition, when studying the link between CEO compensation and firm size, one needs to control for corporate governance and other factors which may affect CEO compensation. One such factor is the *CEO's age*, since it seems reasonable to believe that an executive's age is positively correlated with experience, integrity and skills. Second, *the size of the board* may affect CEO compensation (e.g. Core, Holthausen and Larcker, 1999). For example, a sizeable board can lead to a higher compensation due to the CEO's increased rent-seeking opportunities (e.g. Bebchuk, Fried, and Walker, 2002). Third, *shareholders' ownership concentration* can affect CEO compensation. A large, dominant shareholder may monitor a CEO's actions more effectively, i.e. mitigate potential agency costs, compared to a situation where a firm's ownership is dispersed widely among several shareholders. Therefore, the presence of a dominant shareholder implies lower compensation (e.g. Shleifer and Vishny, 1986). Fourth, *foreign ownership* may affect CEO compensation.[32] For example, foreign investors were perhaps more familiar with option schemes than Finnish shareholders in the past, imposing options on Finnish firms (e.g. Jones, Kalmi and Mäkinen, 2006). As a second model, we pool the data and estimate the following specification:

$$\ln(CEO\ pay_{it}) = \alpha + \beta_1 \ln(Sales_{it}) + \beta_2 CEO\ \mathrm{Age}_{it} + \beta_3 Size\ of\ board_{it} +$$

$$(2) \qquad \beta_4 Dominant\ shareholder(\%)_{it} + \beta_5 Foreign\ ownership(\%)_{it} + \beta_6 ROA(\%)$$

$$\beta_7 Industry\ dummies + \beta_8 Year\ dummies + \varepsilon_{it},\ \mathrm{i=1,...,N},\ \varepsilon_i \sim iid(0,\sigma^2).$$

Our key interest in Equation (2) is on the point estimate of β_1. As a control for firm performance, we use the percentage of ROA (Return on Assets). To control for a possible industry-specific variation in CEO pay, we use three industry dummies: ICT, manufacturing and service. We also add time dummies to control for effects that are common to all firms in a given year. Since the previous empirical studies have used both contemporaneous and once lagged sales as a proxy for firm size, we estimate both contemporaneous and lagged specifications for Equation (2). As in Equation (1), we estimate Equation (2) separately for CEO salary and bonus, and for total compensation. To control for possible omitted variable inconsistency, as a third model we estimate the fixed effects estimator for Equation (2).[33]

Besides focusing on CEO pay-for-firm size elasticity, the literature has explored intensively the relationship between CEO pay and firm performance. According to Conyon and Leech (1994), the Principal-Agent model gives at least a partial theoretical justification for using linear models in this context. Therefore, we next describe briefly the classical Principal-Agent model, where executive compensation is understood as a mechanism to align monetary interests between risk neutral shareholders and risk-averse executives.[34] For example, Holmström and Milgrom (1987) demonstrate that the optimal managerial contract is linear under the assumptions of absence of income effects in an exponential utility function, and an independent and identical distributed error term. The agent's total compensation $\tilde{W}$ includes a constant base salary α and the share β of stochastic output $\tilde{X}$, i.e. $\tilde{W} = \alpha + \beta\tilde{X}$. The power of incentives, i.e. an incentive coefficient β, is decreasing with respect to uncertainty[35], the agent's risk aversion and the agent's effort. Under some restrictive assumptions[36], it is possible to derive the optimal sharing rule

$$(3) \qquad \beta^* = \frac{1}{1 + r\sigma^2 c(e)''}, \text{ where}$$

r is the agent's absolute risk aversion, σ^2 is variance of output (uncertainty) and $c(e)$ is the agent's convex disutility of effort, i.e. $c' > 0, c'' \geq 0$. The optimal sharing rule β* is one, when output $\tilde{X}$ is certain (i.e. $\sigma^2 = 0$) or the agent is risk-neutral (i.e. $r = 0$). Under these circumstances the principal should sell a firm to the agent, which gives the agent maximum monetary incentives. However, when output $\tilde{X}$ is uncertain or the agent is risk-averse, Equation (3) implies that the optimal linear sharing rule β* should be positive in the $0<\beta^*< 1$ range.

In the literature on *CEO pay-for-firm performance sensitivity*, a commonly estimated specification is in first-differences. Thus, instead of estimating models in terms of levels, we focus on the growth of CEO pay, and estimate the following pooled OLS-estimator in terms of first-differences:

$$\Delta \ln(CEO\ pay_{it}) = \alpha + \beta_1 \Delta \ln(Firm\ performance)_{it} + \beta_2 \Delta Size\ of\ board_{it} +$$

$$(4) \qquad \beta_4 \Delta Dominant\ shareholder(\%)_{it} + \beta_5 \Delta Foreign\ ownership(\%)_{it} +$$

$$\beta_6 Industry\ dummies + \beta_7 Year\ dummies + \varepsilon_{it},\ \text{i=1,...,N},\ \varepsilon_i \sim iid(0, \sigma^2).$$

In Equation (4) the dependent variables are the growth of salary and bonuses, and separately, the growth of total compensation. As explained in Section 4.1, we use shareholder wealth and stock return as the measures for firm market-based and the percentage of ROA as an accounting-based performance measures. By doing this, we follow e.g. Lambert and Lackert (1987), who estimated pay-for-firm performance models using both accounting and stock market-based firm performance measures. Finally, some researchers focus on a contemporaneous relationship between CEO pay and firm performance (e.g. Conyon and Murphy, 2000), whereas others use firm performance in period t and in the previous period t-1 (e.g. Hall and Liebman, 1998).[37] Therefore, we also estimate Equation (4) by using both contemporaneous and once lagged firm performance measures.

5. Estimation Results

CEO pay-for-firm size results

Tables 4-6 present the elasticity of CEO compensation with respect to firm size. Table 4 reports cross-section estimates over the period 1996-2002. The estimated elasticity coefficients for CEO cash compensation, i.e. salary and bonuses, with respect to firm sales are all statistically highly significant at the 1% level.[38] The coefficients are in the 0.26-0.34 range supporting the findings of e.g. Baker, Jensen and Murphy (1988) and Rosen (1990) that estimated elasticities do not differ remarkably from 0.3. When focusing on total compensation, the elasticity estimates are in the range of 0.25-0.38. There seems to be a moderate increase in the CEO pay-for-firm size elasticity estimates over time: in the period 1996-1998 the point elasticity estimate for salary and bonuses was 0.29, whereas in 1999-2002 it was 0.32. Similar developments can be noticed for

CEO total compensation, where the point elasticity estimate was 0.32 in 1996-1998, whereas it was 0.36 in 1999-2002.[39] The peak years are 1999 (0.38) and 2000 (0.36).

When comparing our findings to that of Conyon and Murphy (2000), Table 4 suggests that in 1997 a 10% rise in firm sales increased, ceteris paribus, CEO cash compensation approximately 3.2% in the U.S., 2.9% in Finland, and 2.0% in the U.K. Similarly, ceteris paribus, a 10% rise in sales increased CEO total compensation 4.1% in the U.S., 3.0% in Finland and 2.2% in the U.K. The finding suggests that CEO pay-for-firm size elasticity was higher in Finland than in the U.K. but smaller than in the U.S. in 1997, keeping in mind that the number of observations differs between the studies.

[Table 4 about here]

Table 5 shows CEO pay-for-firm size elasticity estimation results for Equation (3), when standard errors are adjust for intragroup correlation. Contrary to Table 4, we now use the pooled OLS estimator and control for CEO age, foreign ownership, ownership concentration, the size of the board, ROA(%), and industry of the firm. In columns (1)-(6) we use sales in period *t* and in columns (7)-(12) in period *t-1* as proxies for firm size. The key finding is that CEO pay-for-firm size elasticity point estimates are close to the range of 0.2-0.3.

There are also some other interesting findings in Table 5. For example, in line with our prior expectations, we find statistically significant evidence that CEO age, foreign ownership, the size of the board, and ROA(%) are positively related to the compensation level. The positive effect of age, approximately 1.0-1.6%, seemed to be transmitted to the base salary and bonus rather than to total compensation. On the contrary, foreign ownership affects both base salary and bonus as well as total compensation. The effect on the former is 0.2-0.4% and on the latter 0.7-1.0%. The effect of ROA% on compensation is moderate, being in the range of 0.5-0.7% for base salary and bonus and in the range of 1.3-1.7% for total compensation. The findings for the size of the board indicate about 7% effect on compensation levels. The presence of a dominant shareholder has about a 0.5% negative effect on compensation levels, being also in the line of prior expectations.

[Table 5 about here]

To control for unobservable fixed effects, we re-estimate Equation (3) using the fixed effects estimator. The estimation results are presented in Table 6. In columns (1)-(6) of Table 6 the firm size elasticity estimates are in the range of 0.26-0.46. When using sales in the period t-1 as a proxy for firm size, the elasticity estimates are in the range of 0.21-0.36. However, in our preferred models in columns (5) and (6), where we have controlled for foreign ownership, dominant shareholder, the size of board and ROA%, the elasticity estimates are 0.3 for base salary and bonus, and 0.46 for total compensation.[40]

We again find statistical evidence that foreign ownership is significantly and positively related to CEO compensation, especially to total compensation. For base salary and bonus the effect is in the range of 0.3-0.5%, whereas for total compensation it is in the 0.9-1.6% range. These numbers are close to those of reported in Table 5. Also, the firm performance measure ROA(%) is positively associated with compensation, and the estimates are close to those of reported in Table 5, indicating a relative robust finding. However, contrary to Table 5, the effect of the size of the board remains significant only for salary and bonus. The estimate indicates about 4% effect of the size of the board on CEO salary and bonus, being about half of that found in Table 5. Ownership concentration is statistically insignificant.[41]

[Table 6 about here]

CEO pay-for-firm performance results

Table 7 presents estimation results for CEO pay-for-firm performance sensitivity, i.e. the model presented in Equation (4). The estimator is the pooled OLS estimator in first-differences, and standard errors are adjust for intragroup correlation. In columns (1)-(6) we use contemporaneous performance measures and in columns (7)-(12) once lagged measures, i.e. firm performance in the previous year.

We find clear statistical evidence on the contemporaneous link between CEO compensation (especially total compensation) and firm stock market performance. For example, our shareholder wealth measures in columns (1) and (2) suggest that the wealth change in CEO salary and bonus is €6.84 per €1,000 change in shareholder wealth. Respectively, the change in CEO total compensation is €21.85 per €1,000 change in shareholder wealth.[42]

In columns (3) and (4), where we use the stock market return as a measure for firm performance, we find that CEO compensation is positively and significantly associated with firm stock return. The parameter estimates are 0.09 for CEO salary and bonus, and 0.28 for total compensation. The estimate 0.09 is close to the range of 0.10-0.15 which Rosen (1990) reports in his survey of several independent studies on CEO compensation. On the contrary, we did not find statistically significant evidence on the association between CEO compensation and firm accounting-based performance in columns (5) and (6).

When using one-year lagged accounting and stock market-based measures for firm performance in columns (8), (10) and (12), we find that it is CEO total compensation that is positively and significantly associated with firm performance. In sum, the empirical findings in Table 7 indicate a significant contemporaneous link between CEO compensation and firm stock market performance, but not with accounting-based performance. However, when using one-year lagged performance measures, we find that only CEO total compensation is positively associated with firm performance. It would be tempting to argue that causality goes from CEO compensation to firm performance. Unfortunately, we cannot answer this question with the current data. It is naturally possible that the causality affects from compensation to performance, but the direction may as well go vice versa. Therefore, we need to be cautious with causality interpretations.

There are also some other interesting findings in Table 7. We find some evidence in columns (5), (6), (10) and (12) that foreign ownership can be positively associated with CEO compensation. Also, in columns (7) and (8), our findings indicate that firm ownership concentration can be negatively related to CEO compensation. Unfortunately, these findings appear to be sensitive to a model specification.

[Table 7 about here]

6. Conclusions

This paper studies how CEO pay is related to firm performance and to firm size in Finland between 1996 and 2002. We utilize new hand-collected individual-level CEO compensation data from the Finnish tax authorities' registers. By providing new empirical evidence from a very different institutional context than the U.S. and the U.K., we hope to increase our understanding on CEO compensation practices across different countries.

When comparing CEO average pay levels to that of average worker compensation in different countries, the evidence suggests that the ratio is substantially higher in the U.S. than in Europe. For example, the ratio was 475 in the U.S. in 1999, whereas in Europe it ranged from 11 in Switzerland to 24 in the U.K. Based on our CEO compensation data, the estimated ratio in Finland (16) in 1999 exceeds the ratio reported in Germany (13), Sweden (13) and France (15), but it is somewhat smaller than in the Netherlands (17), Spain (17) and Belgium (18). When focusing on the recent dynamics of the ratio in Finland, the estimated ratio was 7 in 1996, but peaked at 24 in 2000, likely due to a few executives' exercised stock options during the stock market boom. Thereafter the ratio has fallen to 13 in parallel with the stock market bust in 2000-2002.

It clearly emerges from the data that the level of average CEO compensation increased considerably between 1996 and 2002 in Finland. For example, CEO mean salary and bonus (in real terms) was €165,878 (median €147,113) in 1996, whereas it was €279,733 in 2002 (median €207,856). The percentage increase from 1996 to 2002 is 69% (median 41%). Respectively, CEO mean total compensation increased from €180,190 to €356,863 (98%), whereas median total compensation increased from €155,142 to €232,750 (50%). Since the number of CEOs per year varies in the data, we also calculated the percentage change of the cohort 1996 (from 1996 to 2002). The increase was 74% for both mean and median salary and bonus, and, respectively, 105% (median 74%) for total compensation. In addition, the change in CEO compensation differs notably from that of industrial worker total pay change (11%) from 1996 to 2002.

When focusing on CEO pay-for-firm performance estimates, we find clear statistical evidence that CEO compensation, and especially total compensation, is significantly associated with firm stock market performance. For example, our shareholder wealth measure, which is close to that of Jensen and Murphy (1990), suggests that the contemporaneous change in CEO salary and bonus is €6.84 per €1,000 change in shareholder wealth. Respectively, the change in CEO total compensation is €21.85 per €1,000 change in shareholder wealth. The estimated change in CEO total compensation, i.e. €21.85, is likely to be biased upwards, due to a few large stock option exercises.

Our second stock market-based measure for firm performance, annual stock return, corroborates the previous findings: the change in CEO compensation can be associated with the change in firm performance. For example, for CEO salary and bonus the elasticity estimate is 0.09, and for total compensation it is 0.28. However, when using accounting-based firm performance measure, i.e. ROA%, we do not find statistically significant evidence of the link between compensation and firm performance. However, when using one-year lagged accounting and stock market-based measures for firm performance, we find that total compensation is significantly and positively associated with these performance measures.

Turning to CEO pay-for-firm size elasticity estimates, the estimates do not differ substantially from 0.3, after controlling for CEO age, industry of the firm, ROA%, and three corporate governance indicators (the size of the board, foreign ownership and ownership concentration). An interesting finding is that the elasticity parameter estimates for firm size (proxied by firm sales) are considerably larger than estimates for other explanatory variables. Thus, although the compensation level is positively associated with CEO age, ROA%, foreign ownership and the size of the board, it seems to be firm size that is a key factor in explaining the CEO compensation level. This being the case, our finding supports the CEO compensation-based view for sometimes very active mergers and acquisitions booms, as suggested by Baker, Jensen and Murphy (1988), although naturally other important factors may substantially affect merger and acquisition decisions.

There are also some interesting corporate governance findings. First, the share of foreign ownership is positively and significantly associated with the CEO compensation level. Also, in the most specifications, the foreign ownership parameter estimate is about three times larger for total compensation than for salary and bonus. One possible reason may be that foreign investors are more familiar with equity-based incentives, such as stock options, than Finnish investors. Second, ownership concentration is negatively related to the CEO pay level in the pooled model, supporting the view that a large shareholder can play a substantial role in monitoring CEOs activities. Third, the size of the board is positively related to the CEO pay level, especially to the level of base salary and bonus. This may indicate potential inefficiency, rent-seeking and free-rider issues that can be associated with the functioning of a sizable board.

Tables

Table 1. CEO compensation data summary (in real 2000 Euros).

	1996	1997	1998	1999	2000	2001	2002	Change from 1996 to 2002	2) Change of cohort 1996 from 1996 to 2002
# of CEOs (% of listed firms total)[1]	43 (58.9%)	54 (47.0%)	62 (52.1%)	74 (54.0%)	82 (54.7%)	78 (53.8%)	71 (51.1%)	+65%	
CEO mean salary + bonus, € (% change from previous year)	165,878	174,288 (+5.1%)	204,679 (+17.4%)	203,233 (-0.7%)	220,549 (+8.5%)	257,818 (+16.9%)	279,733 (+8.5%)	+69%	+74%
CEO median salary + bonus, € (% change from previous year)	147,113	174,337 (+18.5%)	194,582 (+11.6%)	176,353 (-9.4%)	193,495 (+9.7%)	199,250 (+3.0%)	207,856 (+4.3%)	+41%	+74%
CEO mean total compensation, € (% change from previous year)	180,190	205,481 (+14.0%)	318,865 (+55.2%)	422,974 (+32.7%)	646,845 (+53.0%)	375,439 (-42.0%)	356,863 (-4.9%)	+98%	+105%
CEO median total compensation, € (% change from previous year)	155,142	183,434 (+18.2%)	206,736 (+12.7%)	202,542 (-2.0%)	205,891 (+1.7%)	218,384 (+6.1%)	232,750 (+6.6%)	+50%	+74%
Industrial worker mean total compensation, € (% change from previous year)	24,793	25,000 (+0.8%)	25,379 (+1.5%)	25,973 (+2.3%)	26,435 (+1.8%)	27,130 (2.6%)	27,660 (+2.0%)	+12%	
Ratio between CEO and industrial worker mean total compensation	7	8	13	16	24	14	13		
Mean EBIT, €1000 (% change from previous year)	48,752	91,078 (+86.8%)	109,800 (20.6%)	116,734 (+6.3%)	150,000 (+28.5%)	123,493 (-17.7%)	140,721 (+14.0%)	+189%	
Median EBIT, €1000 (% change from previous year)	16,215	18,770 (+15.8%)	17,650 (-6.0%)	16,839 (-4.6%)	13,800 (-18.0%)	17,406 (+26.1%)	16,578 (-4.8%)	+2%	
HEX portfolio index (Log change from previous year, %)	-	+27.3%	+13.8%	+54.1%	-24.2%	-19.1%	-15.0%		
Mean age	47.6	48.6	49.3	50.1	50.8	51.6	52.2		
Median age	48.5	50.0	50.5	50.0	51.0	52.0	52.5		
# of CEOs, who exercised stock options (% of CEOs in data)	1 (2.3%)	3 (5.6%)	8 (12.9%)	9 (12.2%)	9 (11.0%)	10 (12.8%)	11 (15.7%)		
Mean value of exercised stock options, € (exercised options as a % of total compensation)	173,283 (2.2%)	369,137 (10.0%)	793,856 (32.1%)	1,714,511 (49.2%)	3,787,800 (64.2%)	919,949 (28.3%)	966,671 (42.5%)		

[1] We have information of all listed firms from 1997 onwards. Thus, for 1996 the number of listed firms is firms in the main list. Note that the CEO wage distributions are skewed to the right, i.e. the means are typically higher than the medians. As a deflator we have used the GDP deflator obtained from Statistics Finland.

[2] In the column we compare the percentage change in the compensation of 32 CEOs of the 1996 cohort for whom we have compensation information over the whole period 1996-2002.

Table 2. Summary statistics for year 2000.

Variable	N	Mean	Median	Standard deviation
CEO salary and bonus, €	82	220,549	193,495	159,647
CEO total compensation, €	82	646,845	205,891	2,538,019
CEO age	82	50.75	51	5.98
Firm sales, €	82	1,370,000,000	251,000,000	3,820,000,000
Size of board	77	6.12	6	1.65
EBIT, €	82	150,000,000	13,800,000	680,000,000
ROA, %	81	10.92	9.7	10.82
Foreign ownership, %	82	20.48	16.5	20.46
Dominant shareholder's ownership, %	82	22.14	19.04	15.33
Annual stock return, %	82	-18.36	-8.51	52.16

Table 3. CEO total compensation in Finland, Norway and Sweden in 1998.

	Finland	Norway	Sweden
# of firms in data sample	62	120	104
Mean total compensation, €	310,732	161,670	279,249
Median total compensation, €	189,619	N/A	N/A
% of firms that paid a bonus	N/A	N/A	42% (44 firms)
Mean value of a bonus, € (bonuses as a % of total compensation)	N/A	N/A	42,057 (23%)
% of CEOs that exercised options	13% (8 CEOs)	N/A	N/A
Mean value of exercised stock options, € (exercised options as a % of total compensation)	773,609 (32%)	N/A	N/A
Mean total compensation without those 8 CEOs that exercised stock options, €	185,308	-	-

N/A= not available. CEO compensation data for Norway and Sweden is from Randøy and Nielsen (2002). CEO mean total compensation includes all pay components.

Table 4. CEO pay-for-firm size elasticity in the U.S., U.K. and Finland.

	Elasticity of CEO pay with respect to firm sales	# of CEOs
U.S., 1997		
Salary and bonus	0.316	1,666
Total pay	0.413	1,666
U.K., 1997		
Salary and bonus	0.197	510
Total pay	0.217	510
FINLAND, 1996		
Salary and bonus	0.263	43
Total pay	0.249	43
FINLAND, 1997		
Salary and bonus	0.291	54
Total pay	0.299	54
FINLAND, 1998		
Salary and bonus	0.286	62
Total pay	0.335	62
FINLAND, 1999		
Salary and bonus	0.301	74
Total pay	0.376	74
FINLAND, 2000		
Salary and bonus	0.312	82
Total pay	0.364	82
FINLAND, 2001		
Salary and bonus	0.335	78
Total pay	0.350	78
FINLAND, 2002		
Salary and bonus	0.326	71
Total pay	0.330	71
FINLAND, 1996-1998		
Salary and bonus	0.285	159
Total pay	0.305	159
FINLAND, 1999-2002		
Salary and bonus	0.321	305
Total pay	0.357	305

1. The estimated model is $\ln(\text{CEO salary and bonus}_i)=\alpha+\beta\ln(\text{firm sales}_i) + \varepsilon_i$.
2. All estimated coefficients for Finnish CEOs are statistically significant at the 1% level, based on robust standard errors (i.e. Huber-White-sandwich estimator of variance). The Breusch-Pagan test implied heteroskedasticity in salary and bonus estimations for years 2001 and 2002, and in total pay estimations for years 1998-2000.
3. The elasticity estimates for the U.S. and the U.K. are from Conyon and Murphy (2000).

Table 5. CEO pay-for-firm size elasticity (the pooled OLS estimator) 1996-2002.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)
Constant	5.136 ***	4.507 ***	5.650 ***	5.596 ***	5.756 ***	5.436 ***	5.355 ***	5.417 ***	5.799 ***	5.849 ***	6.067 ***	5.805 ***
	(0.248)	(0.381)	(0.284)	(0.392)	(0.268)	(0.429)	(0.248)	(0.308)	(0.284)	(0.409)	(0.278)	(0.460)
ln (sales)$_t$	0.308 ***	0.363 ***	0.285 ***	0.299 ***	0.247 ***	0.270 ***	-	-	-	-	-	-
	(0.010)	(0.022)	(0.013)	(0.021)	(0.012)	(0.025)						
ln (sales) $_{t-1}$	-	-	-	-	-	-	0.311 ***	0.372 ***	0.283 ***	0.295 ***	0.246 ***	0.269 ***
							(0.010)	(0.024)	(0.013)	(0.023)	(0.013)	(0.029)
ROA, %	0.005 ***	0.013 ***	0.004 **	0.015 ***	0.006 ***	0.015 ***	0.006 ***	0.014 ***	0.006 ***	0.016 ***	0.007 ***	0.017 ***
	(0.002)	(0.004)	(0.002)	(0.004)	(0.002)	(0.004)	(0.002)	(0.004)	(0.002)	(0.004)	(0.002)	(0.004)
CEO age	0.014 ***	0.005	0.016 ***	0.009 *	0.016 ***	0.012 **	0.010 ***	-0.001	0.012 ***	0.004	0.013 ***	0.006
	(0.004)	(0.005)	(0.004)	(0.006)	(0.004)	(0.006)	(0.004)	(0.006)	(0.004)	(0.006)	(0.004)	(0.007)
Foreign ownership, %	-	-	0.003 ***	0.009 ***	0.002 *	0.007 ***	-	-	0.004 ***	0.010 ***	0.003 **	0.009***
			(0.001)	(0.002)	(0.001)	(0.002)			(0.001)	(0.002)	(0.001)	(0.002)
Dominant shareholder's ownership, %	-	-	-0.004 ***	-0.006 ***	-0.004 ***	-0.006 ***	-	-	-0.003 ***	-0.006 ***	-0.004 ***	-0.006 ***
			(0.001)	(0.001)	(0.001)	(0.001)			(0.001)	(0.002)	(0.001)	(0.002)
Board size	-	-	-	-	0.076 ***	0.071 ***	-	-	-	-	0.065 ***	0.068 ***
					(0.013)	(0.022)					(0.014)	(0.026)
Manufacturing	-	-	-	-	0.059	0.109 **	-	-	-	-	0.068	0.129 **
					(0.040)	(0.052)					(0.044)	(0.060)
ICT	-	-	-	-	-0.059	0.131	-	-	-	-	-0.044	0.170 *
					(0.054)	(0.088)					(0.055)	(0.094)
R^2	0.75	0.61	0.71	0.60	0.76	0.62	0.75	0.59	0.73	0.60	0.76	0.61
Firms	87	87	86	86	82	82	87	87	86	86	79	79
Observations	453	453	415	415	377	377	376	376	354	354	318	318

1. The dependent variable is in natural logarithms. All monetary variables are deflated by an industry-specific gross-output deflator in 2000 Euros.
2. Standard errors are adjusted for intragroup correlation in the parentheses. ***, **, * statistically significant at 1%, 5% and 10% levels, respectively.
3. All models include a full set of year dummies. Service sector is a reference group for industry dummies.

Table 6. CEO pay-for-firm size elasticity (the fixed effects estimator) 1996-2002.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)	ln(salary and bonus)	ln(total pay)
ln (sales)$_t$	0.266 *** (0.038)	0.304 *** (0.066)	0.266 *** (0.043)	0.361 *** (0.097)	0.323 *** (0.050)	0.503 *** (0.125)	-	-	-	-	-	-
ln (sales) $_{t-1}$	-	-	-	-	-	-	0.237 *** (0.038)	0.264 *** (0.084)	0.179 *** (0.045)	0.194 ** (0.100)	0.248 *** (0.057)	0.316 *** (0.129)
Foreign ownership, %	-	-	0.003 * (0.002)	0.010 *** (0.004)	0.003 * (0.002)	0.011 *** (0.004)	-	-	0.003 * (0.002)	0.010 *** (0.004)	0.006 *** (0.002)	0.017 *** (0.005)
Dominant shareholder's ownership, %	-	-	-0.001 (0.002)	0.002 (0.003)	-0.001 (0.002)	0.004 (0.003)	-	-	-0.001 (0.002)	0.000 (0.004)	-0.000 (0.002)	0.004 (0.005)
CEO age	-	-	0.035 *** (0.006)	0.050 *** (0.010)	0.033 *** (0.005)	0.054 *** (0.011)	-	-	0.032 *** (0.007)	0.047 *** (0.017)	0.032 *** (0.007)	0.052 *** (0.016)
Board size	-	-	-	-	0.030 ** (0.015)	-0.009 (0.030)	-	-	-	-	0.030 ** (0.017)	-0.000 (0.037)
ROA, %	-	-	-	-	0.009 *** (0.002)	0.017 *** (0.005)			-	-	0.009 *** (0.003)	0.017 *** (0.006)
R^2 (within)	0.51	0.30	0.52	0.32	0.56	0.36	0.49	0.20	0.49	0.22	0.52	0.27
Firms	87	87	86	86	82	82	87	87	86	86	79	79
Observations	464	464	422	422	377	377	377	377	355	355	318	318

1. The dependent variable is in natural logarithms. All monetary variables are deflated by an industry-specific gross-output deflator in 2000 Euros.
2. Robust standard errors in the parentheses. ***, **, * statistically significant at 1%, 5% and 10% levels, respectively.
3. All models include a full set of year dummies.

Table 7. CEO pay-for-firm performance sensitivity (the pooled OLS estimator in first-differences).

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	Δln(salary and bonus) POLS 1996-2002	Δln(total pay) POLS 1996-2002	Δln(salary and bonus) POLS 1996-2002	Δln(total pay) POLS 1996-2002	Δln(salary and bonus) POLS 1996-2002	Δln(total pay) POLS 1996-2002	Δln(salary and bonus) POLS 1996-2002	Δln(total pay) POLS 1996-2002	Δln(salary and bonus) POLS 1996-2002	Δln(total pay) POLS 1996-2002	Δln(salary and bonus) POLS 1996-2002	Δln(total pay) POLS 1996-2002
Constant	0.090 **	0.200 **	0.092 **	0.207	0.039	-0.012	0.070	0.166	0.032	-0.019	0.091 *	0.186
	(0.040)	(0.087)	(0.040)	(0.088)	(0.069)	(0.082)	(0.062)	(0.122)	(0.077)	(0.092)	(0.049)	(0.105)
Δ(shareholder wealth)$_t$	0.00684 ***	0.02185 ***	-	-	-	-	-	-	-	-	-	-
	(0.003)	(0.006)										
Δ(shareholder wealth)$_{t-1}$	-	-	-	-	-	-	0.00348	0.01533 **	-	-	-	-
							(0.003)	(0.007)				
Δ(stock return)$_t$	-	-	0.0931 **	0.2781 ***	-	-	-	-	-	-	-	-
			(0.037)	(0.082)								
Δ(stock return)$_{t-1}$	-	-	-	-	-	-	-	-	0.0480	0.1828 **	-	-
									(0.039)	(0.082)		
Δ(ROA)$_t$	-	-	-	-	0.001	0.010	-	-	-	-	-	-
					(0.002)	(0.010)						
Δ(ROA)$_{t-1}$	-	-	-	-	-	-	-	-	-	-	0.002	0.015 *
											(0.004)	(0.008)
Δ(foreign ownership), %	0.002	0.008	0.002	0.008	0.004 *	0.014 ***	0.001	0.009	0.003	0.010 *	0.005	0.017 **
	(0.002)	(0.005)	(0.002)	(0.005)	(0.002)	(0.005)	(0.003)	(0.006)	(0.002)	(0.005)	(0.004)	(0.008)
Δ(dominant shareholder's ownership), %	-0.002	-0.001	-0.002	-0.001	-0.006	-0.006	-0.011 **	-0.013 *	-0.006	-0.004	-0.010 *	-0.008
	(0.003)	(0.005)	(0.003)	(0.005)	(0.004)	(0.005)	(0.005)	(0.007)	(0.004)	(0.005)	(0.005)	(0.008)
Δ(board size)	0.010	-0.024	0.010	-0.024	0.012	-0.021	-0.022	-0.049 *	0.011	-0.021	-0.002	-0.023
	(0.020)	(0.026)	(0.020)	(0.027)	(0.021)	(0.028)	(0.018)	(0.028)	(0.021)	(0.028)	(0.020)	(0.029)
Manufacturing	0.054*	0.095 *	0.055 *	0.097 *	0.050	0.069	0.033	0.063	0.040	0.075	0.020	0.047
	(0.031)	(0.051)	(0.031)	(0.052)	(0.033)	(0.052)	(0.034)	(0.063)	(0.032)	(0.054)	(0.031)	(0.057)
ICT	0.079 *	0.221 *	0.076 *	0.208 *	0.044	0.179	0.014	0.109	0.037	0.128	0.048	0.191
	(0.044)	(0.126)	(0.044)	(0.125)	(0.047)	(0.117)	(0.034)	(0.160)	(0.045)	(0.120)	(0.052)	(0.164)
R^2	0.05	0.11	0.05	0.10	0.06	0.09	0.11	0.10	0.05	0.08	0.11	0.12
Firms	77	77	77	77	77	77	65	65	77	77	70	70
Observations	301	301	301	301	294	294	225	225	299	299	231	231

1. All monetary variables are deflated by an industry-specific gross-output deflator in 2000 Euros.
2. All models include a full set of year dummies.
3. Standard errors are adjusted for intragroup correlation in the parentheses. ***, **, * statistically significant at 1%, 5% and 10% levels, respectively.
4. The service sector is a reference group for industry dummies.

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Endnotes

[1] I am extremely indebted to Finnish tax authorities for providing access to individual level CEO compensation data and to Balance Consulting for firms' financial statement data. I gratefully acknowledge the LIIKE research grant from the Academy of Finland. In addition, I thank the Yrjö Jahnsson Foundation, the Helsinki School of Economics Research Foundation, the Foundation of Kluuvi, and the Marcus Wallenberg Foundation for financial support. All support from the Research Institute of the Finnish Economy (ETLA) is gratefully acknowledged. I thank Pekka Ilmakunnas, Panu Kalmi and Otto Toivanen for their helpful comments and suggestions. An earlier version of this paper was presented at the annual meeting of Finnish Economists in Oulu, February 2003, at the Compensation session of 30th conference of the EARIE in Helsinki, August 2003, and at the EIASM workshop on Corporate Governance in Brussels, November 2005, and at the Corporate Governance session of 33rd conference of the EARIE in Amsterdam, August 2006. We thank all the participants for their helpful comments. The usual disclaimer applies.

[2] In a modern company, which saw daylight in the industrial revolution, ownership is usually separated from control. As a consequence, executives have more information than shareholders about a firm's possible risks and returns.

[3] Fama (1980), and Fama and Jensen (1983) have also emphasised the principal-agent problem and the separation of ownership from control.

[4] The size of the board, the voting share of a largest shareholder, and the share of foreign ownership have been used as proxies for corporate governance in the literature.

[5] Jensen (1993) outlines four categories for corporate governance mechanisms: legal and regulatory, internal, external and product market competition. Since these categories are not perfectly distinct, as he also underlines, we use only two categories: internal and external mechanisms, where external mechanisms include both legal and regulatory and product market competition mechanisms. For more comprehensive discussion of corporate governance mechanisms, see e.g. Monks and Minow (2001).

[6] As a company form, a modern public limited company substantially reduces uncertainty of sharing returns and risks among stakeholders. For example, corporate legislation determines how firm profits should be shared among stakeholders. However, despite carefully enacted corporate legislation, a modern firm cannot remove all the uncertainty between stakeholders. The principal-agent problem between shareholders and executives will always exist due to asymmetric information.

[7] The Helsinki Stock Exchange, the Central Chamber of Commerce of Finland and the Confederation of Finnish Industry and Employers appointed a working group on February 2003 to clarify the need of reviewing the corporate governance recommendation that was issued by the Central Chamber of Commerce of Finland and the Confederation of Finnish Industry and Employers in 1997. The working group published the revised recommendations on December 2003. These recommendations can be found in http://www.keskuskauppakamari.fi/kkk/en_GB/etusivu/ (13.3.2006).

[8] This should not be confused with the concept of maximising shareholders profits in the short run, since then a firm's total value will not be maximised. As argued by Jensen, Murphy and Wruck (2004), the objective for firm total value maximising says that e.g. employees' satisfaction and product quality should be increased to a point where a future marginal increase in each reduces firm value.

[9] Therefore, from a risk management viewpoint, it appears to be important that the best codes of corporate governance practices are implemented if executives are awarded equity-based incentives. Naturally, there is no executive compensation policy that "equally fits for all firms". Instead, coherent executive remuneration policy requires both an understanding of a firm's strategy, goals, and vision, and knowledge of feasible compensation possibilities and corporate governance practices.

[10] For a more detailed exposition of law and financial changes related to corporate governance in Finland, see e.g. Hyytinen, Kuosa and Takalo (2003).

[11] See e.g. Kiander and Vartia (1996), and Honkapohja and Koskela (1999).

[12] See e.g. Jones, Kalmi and Mäkinen (2006).

[13] See e.g. Piekkola (2005).

[14] See e.g. Ylä-Anttila, Ali-Yrkkö and Nyberg (2004).

[15] Murphy (1999) provides empirical support for the key role of stock options: "... our analysis shows that pay-performance sensitivity has nearly doubled to \$6.0 per \$1,000 change in shareholder value by 1996. The increase in pay-performance sensitivities has been driven almost exclusively by stock option grants."

[16] For example, in the U.S. it is compulsory for publicly listed firms to disclose information on the top five executives' compensation. Also, nowadays the trend in corporate governance regulations in other

countries is more often to recommend that publicly listed firms disclose detailed information on CEO compensation.

[17] In addition, findings in the previous literature are based on CEO compensation data from publicly listed firms.

[18] To obtain individual CEO annual compensation data from the Finnish tax authorities' registers, we needed to have a social security number for each CEO. All possible identity numbers were hand-collected from different sources, such as from the National Board of Patents and Registration of Finland's public registers. Though our CEO compensation data set is not a random sample from the population of listed firms' CEOs, we believe it fairly well presents the compensation pattern of Finnish CEOs in the listed firms, since the number of our CEOs encompasses around 50% of listed firms' CEOs.

[19] A description of these handbooks may found in http://www.porssitieto.fi/index.html (22.3.2005). We thank Iikka Kuosa for providing some sample data he had collected.

[20] Therefore, we are able to separate a CEO's actual compensation obtained from his firm and from other sources.

[21] Typically, a CEO's compensation package may include a company car, a mobile phone, lunch benefits etc. We do not observe individual CEO compensation contracts from our compensation data, but we do perceive the yearly taxation value for these benefits. According to our calculations these benefits vary typically between €10.000 and €20.000 per year. We set the "critical limit value" to be clearly higher, i.e. €35.000 per year over the period 1996-1999, than the typical taxation value. Thus, we believe that the exercised stock options are almost surely the only reason for the values that are greater than €35.000. From 2000 onwards we have observed directly the value of exercised stock options.

[22] Unfortunately, CEO compensation data does not contain any information on granted but not exercised options. Therefore we have to bypass the possible incentive effects of unvested stock options. We also ignore all incentive effects from firm stocks, because we do not have information on CEOs' stock ownership. However, typically a Finnish CEO owns a quite moderate amount of stocks of his employer. This complicates comparison to Jensen and Murphy (1990), who used the information of stock ownership, unvested options etc. by summing eight CEO pay components in estimating total pay-for-performance sensitivity.

[23] See e.g. Conyon and Murphy (2000).

[24] See e.g. Conyon and Leech (1994).

[25] See e.g. Rosen (1990), and Kato and Kubo (2005).

[26] That is to say the number of CEOs encompasses 47-59% of the firms listed on the Helsinki Stock Exchange (one firm-one CEO), depending on the year.

[27] Information on other European countries is based on The Economist (2000).

[28] Although the U.S. ratio has fallen close to 160 in 2004, there seems still to be a remarkable difference in the ratios between the U.S. and Europe.

[29] The median EBIT in real 2000 euros has also increased, but the percentage change from 1996 to 2002 is only 2%. This might be explained by the skewed EBIT distribution to the right.

[30] Unfortunately, our compensation data do not reveal if a CEO has unexercised stock options. However, a typical exercise pattern in Finland is that CEOs exercise their options very soon after the first possible exercise day.

[31] Compensation data on Norway and Sweden is based on Randøy and Nielsen (2002). The number of CEOs varies between Finland, Norway and Sweden in Table 3, which may exacerbate the comparison of compensation levels.

[32] This measures the percentage of firm shares held by foreign citizens and institutions. Unfortunately, we are unable to identify the home country of foreign shareholders.

[33] The classic example is the ability bias in estimating the effects of education on individual earnings. The fixed effects estimator, however, allows us to control for unobserved time-invariant effects to the extent that their effect on the conditional mean is the same in each year. Second, there is a potential correlation between an observable explanatory variable and an unobservable individual effect. This also supports the use of the fixed effects estimator, since the estimator produces asymptotically consistent parameter estimates regardless of the correlation between an observable explanatory variable and an unobservable individual effect. Note, however, that this robustness has a substantial cost: there needs to be individual and time-variation in an explanatory variable or we cannot distinguish the effect of this variable from the effect a constant unobservable variable.

[34] Theoretically this approach utilises the contract theory and the moral hazard (or hidden action) models highlighting the trade-off between insurance (a fixed wage) and incentives (a variable component in a wage contract). The models typically assume that a principal does not observe an agent's effort.

[35] Prendergast (2002) documents the results of 11 empirical studies for the trade-off of uncertainty and incentives for executives. The empirical results are mixed, i.e. there may be a statistically negative, zero or positive relationship between the increase of uncertainty and pay for performance (incentives).

[36] See e.g. Holmström and Milgrom (1987), Murphy (1999).

[37] The reason is that the bonus part of compensation is typically determined in the end of a year reflecting a contemporaneous link between CEO compensation and firm performance. On the contrary, the reason for using a once lagged firm performance measure is that CEO salary is commonly set at the beginning of the year, being sensitive to firm performance in the previous year.

[38] Standard errors are Huber-White-sandwich heteroskedastic robust estimates. All estimations are made using the STATA/SE 9.1 statistical package.

[39] Note that these findings are indicative, since the elasticity estimates are point estimates.

[40] These are somewhat higher than the estimates based on the pooled OLS estimator in columns (5) and (6) of Table 5, but also the models are different, e.g. we cannot control for industries in Table 6.

[41] One reason might be a modest variation of ownership (and the size of board) over time *t*, though we have plenty of variation over individuals *i*.

[42] The estimated change in CEO total compensation, i.e. €21.85, is likely to be biased upwards, due to a few large stock option exercises.

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