The economic impact of capital expenditures: Environmental regulatory delay as a source of competitive advantage?

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Abstract

This study tests the proposal that by undertaking voluntary capital expenditures that are subject to lengthy environmental regulatory delays, listed companies can gain a competitive advantage. The stock market is found to react positively to new capital expenditure announcements when projects are expected to experience long delays in obtaining environmental regulatory approval. Two sources of potential competitive advantage are firm learning and first mover advantages. Lengthy delays in regulatory processes and high compliance costs incurred for environmentally-sensitive projects may allow firms opportunities to develop specialised capabilities and/or to deter industry competitors and new entrants, resulting in greater expected project NPVs. The findings also underscore the importance of nonfinancial environmental information to investors in their assessment of firm value.

Keywords: Regulatory delay, capital investments, competitive advantage, shareholder wealth creation, market valuation, environmental regulation, compliance costs, nonfinancial information, strategy, New Zealand

JEL Classifications: G31, Q58

1. Introduction

The costs and uncertainties caused by delays to new project approvals that are subject to stringent environmental regulatory processes are of major concern to businesses and their investors (Wood, 2003). While some research has attempted to quantify environmental compliance costs (e.g. Joshi, et al., 2001), the possible firm-level benefits associated with regulatory delays have received scant attention. This paper seeks to address this gap in the literature. The main proposal to be tested is that by undertaking voluntary capital expenditures with high environmental compliance costs, listed companies can gain a strategic advantage. Competitive advantages may be gained if regulatory delays enhance firms' opportunities to develop specialised capabilities and resources, such as early mover advantages, reputational benefits, or sophisticated environmental management systems (Hart, 1995). Further competitive benefits may accrue to firms in the form of greater expected project NPVs if the level of resource consent compliance costs incurred are sufficient to pre-empt actions by industry competitors and new entrants (Dean and Brown, 1995, Ryan, 2005).

The environmental approval processes for major projects in most developed (and some developing) countries tend to be lengthy. Typically, overarching legislation applies at the national level, while implementation is at a state or regional level. For example, in the US, the 1969 National Environmental Policy Act sets the requirements for major projects to undergo an assessment of environmental impacts (EIA), while individual state legislation controls firm permitting, monitoring and enforcement activities. European Union Countries follow the 1985 European Union Directive (85/337/EEC) on EIA while setting specific standards nationally.[1] Comparing EIA systems in seven countries, Wood (2003) notes that complaints of excessive delays in such processes are common, most notably in the US,

[1] For example, in England and Wales, the Town and Country Planning (Environmental Impact Assessment) Regulations 2011 applies.

Canada and Australia. He observes that these jurisdictions have relatively formalised EIA systems that require applicants to undertake numerous procedural steps. Comparing the air pollution emission permitting processes for four automobile assembly plants in the US and Germany, Dwyer et al. (1999) conclude that the extensiveness of public consultation, stringency of permit requirements, and degree of federal oversight are greater in the US. These reasons are suggested to explain the differences in the time to approve the permits, being less than two years in Germany as opposed to several years in the US. Investigating delays in the approval of operating permits in the mid-west US, Decker (2003) finds the average time to process permit applications for new industrial projects depends on the expected environmental sensitivity of the project, being 14 months when sensitivity is low and 30 months when it is high. In contrast, the average time to obtain a license to build hydropower plants over a sample of 214 Norwegian energy projects between 2001 and 2008 was relatively quick at one and a half years (Heggedal, et al., 2011).

New Zealand (NZ) is a case-in-point where the consenting process required by the Resource Management Act (RMA, 1991) has drawn wide-spread criticism for causing excessive delays, uncertainties and compliance costs for businesses (Ministerial Panel on Business Compliance Costs, 2001, OECD, 1996, 2007). The legislation requires individuals and businesses to apply for resource consent approval prior to using resources or undertaking development activities that may have potentially adverse environmental effects. Governance over the use of resources is devolved to regional and local bodies, so multiple consents may be required to allow major investment projects to proceed when more than one jurisdiction or type of resource is affected.[2] Reasons suggested for the processing delays include inconsistency of approval processes (Ministerial Panel on Business Compliance Costs, 2001, OECD, 1996, 2007), excessive risk aversion by authorities (Upton,

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[2] Once granted, resource consents are only transferable as part of the project assets. An exception to this is water rights, but in practice their tradability is limited.

1997) and the presence of nuisance objectors and trade competitors during the consultative stages (Ernst & Young, 1997, Ministerial Panel on Business Compliance Costs, 2001).

Quantifying environmental compliance costs can be difficult, as there appear to be impediments to their disclosure through annual financial reports. Resource consent costs are not required to be separately identified in NZ company reports, and are either expensed if they fail to meet the asset test, or capitalised with the project assets.[3] In the US steel industry, research suggests that the ratio of indirect to direct environmental compliance costs is approximately 10:1, with only direct costs being separately identified in the accounting system (Joshi, et al., 2001).

In contrast, resource consent information is widely disseminated using corporate announcements through the stock exchange and media. From 1 December 2002, the New Zealand Exchange (NZX) continuous disclosure regulations require listed companies to immediately release information that is expected to have a material effect on their share prices (New Zealand Exchange, 2005).[4] Typically, companies first disclose the nature of the proposed project and resource consent plans. Investors may use this information to predict the expected time needed to gain resource consent approval, consider the probability of success/abandonment and assess the project valuation implications which are then impounded into stock prices. Later as delays arise or milestones are reached in the consenting process, companies make further announcements which may have further valuation implications. In their analysis of the informativeness of resource consent disseminations over the course of the consenting process, Wirth et al. (2011) contend that

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[3] Tozer and Hawkes (2001) reveal that out of 24 NZ companies surveyed that held resource consents, only one reported them separately in the financial statements.

[4] Previously, price-sensitive information possessed by a listed company was required to be immediately disclosed to the market when the value of confidentiality no longer exceeded the value of disclosure (Huang, et al., 2009).

the nonfinancial nature of the disclosures makes it difficult for investors to assess the expected project cash flows and risks. Yet prior research documents that nonfinancial information communicated independently from financial statements can have economic value. Examples include third-party patent quality data for high-tech firms (Hirschey, et al., 2001), media-reported customer satisfaction scores (Ittner and Larcker, 1998), third-party customer retention and usage data in the wireless industry (Livne, et al., 2011) and firm web site disclosures of firm value creation indicators (Orens, et al., 2010). In this paper, we present evidence suggesting that investors use nonfinancial information from the stock exchange and media to estimate future compliance costs associated with proposed projects.

Following Porter's (1979) assertion that strategic capital expenditures may provide firms with competitive advantages, this study uses NZ data to test the proposition that projects with higher consent compliance costs may provide listed companies with first-mover or other sustainable advantages that make them more valuable. To do this, the forecast time for each project to obtain resource consent approval is used as an indicator of expected resource consent compliance costs at the time of each project initiation announcement. This forecast indicator is found to be positively related to the actual time taken to either obtain consent approval or abandon the project. Using event study methodology, a positive stock market reaction to a sample of NZ project initiation announcements is documented from 1992 to 2007. We then divide the sample at the median expected time to gain consent approval, and consistent with our proposition, we find that the positive valuation effects are significant only for those projects that are expected to experience longer regulatory delays.

This research contributes to the literature in several ways. First, contrary to the negative attention that consent/permitting processes have gained, we provide evidence of a positive valuation impact of environmental compliance costs that supports previous research suggesting that the wealth of regulated firms may increase when environmental regulations assign property rights to environmental assets (e.g. Maloney and McCormick, 1982).

Second, we take the novel approach of using regulatory delay as an indicator of environmental compliance costs, which other researchers may wish to apply to other jurisdictions. Third, compared with compliance cost surveys that only measure costs, our focus on capital expenditures announcements allows insights into net benefits at the project-level. Fourth, this study directly addresses the criticism by Jaffee et al. (1995) in their literature review of the economic impacts of environmental regulation, that most studies fail to control for costs of delays and litigation caused by environmental regulation. Fifth, the findings offer insights into possible trade-offs faced by regulators with respect to the impact of environmental protection policies on capital investment, wealth creation and market structure. Finally, the results provide evidence that investors use nonfinancial environmental information in conjunction with more conventional financial information in their assessment of firm value.

The paper is structured as described next. Section 2 provides a brief review of prior literature and develops the hypotheses. The data and methodologies employed are discussed in Section 3, and the empirical results are presented in Section 4. Concluding remarks are made in Section 5.

2. Literature review and hypothesis development

The proposition that environmental regulatory delays for new project approvals can provide firms with an economic benefit is derived principally from strategic management and environmental economics literature. According to the strategic management view proposed by Porter (1979), firms may gain competitive advantages by undertaking new, strategic capital expenditures that require substantial financial resources. Porter and van der Linde (1995) develop this idea further in an environmental economics context by suggesting that well-designed environmental regulations may encourage resource productivity, enhance innovation and improve competitiveness. Two sources of possible benefit identified in the literature are firm learning and first mover advantages.

The firm learning argument proposed in Hart's (1995) 'natural-resource-based view of the firm' suggests that valuable internally-developed capabilities and resources within a firm may provide sustainable competitive advantages. This implies that lengthy delays in regulatory processes may allow firms opportunities to gain expertise to develop more sophisticated environmental risk management systems and better manage permitting processes. Superior environmental performance has been found to be positively associated with improved operational performance (Melnyk, et al., 2003), profitability (Russo and Fouts, 1997) and firm value (King and Lenox, 2002). Furthermore, first movers may develop expertise that creates valuable real options (Bernardo and Chowdhry, 2002).

A first mover strategy may be beneficial, given that there is a limit to the number of wind farms that can occupy a mountain range, hydro-generation projects that can use a specific water resource, or retirement villages that can enjoy a particularly strategic location. The first mover advantage literature suggests that early adopters can earn economic rents by investing in cost-reducing technology during the early stages of an industry life cycle (Jovanovic and Macdonald, 1994). Empirical evidence finds that first movers and firms undertaking voluntary environmental capital expenditures derive economic advantages from their investment activities (Johnston, 2005, Nehrt, 1996). Some literature supports the view that the regulatory allocation of property rights through quotas, licenses and permits may restrict industry outputs or deplete common resources, ultimately creating barriers to new industry entrants and economic profits for industry incumbents (Buchanan and Tullock, 1975, Maloney and McCormick, 1982, Mason and Polasky, 1994). Further studies show that sunk costs caused by the irreversibility of environmental capital expenditures may also impose barriers to new firm entry (e.g. Dean and Brown, 1995, Ryan, 2005).

This study uses event study methodology to test the shareholder wealth implications of capital asset expenditures that must comply with environmental regulations. Previous event studies of capital expenditure announcements often test aspects of the investment opportunities theory, which suggests that managers are able to maximise the market value of the firm by undertaking positive NPV projects (Miller and Modigliani, 1961). Some empirical studies find support for shareholder wealth maximisation through positive market reactions to news of increases in capital budgets (McConnell and Muscarella, 1985), particularly for firms with greater investment opportunities (Chung, et al., 1998, Vogt, 1997). Similarly, positive valuation implications have been found in studies of capital expenditure project announcements. Woolridge and Snow (1990) find that strategic investment projects earn positive abnormal returns irrespective of the expected investment horizon, while Jones, Danbolt and Hirst (2004) document that announcements of projects that create growth options are associated with significantly greater market-adjusted returns than announcements of projects that exercise growth-options. In their study of the intra-industry effects of US corporate capital expenditure announcements, Chen et al. (2007) report a positive impact on announcers' event-period abnormal returns and a net negative effect on competitors' market values, which they consider to be consistent with the competitive advantages suggested by Porter (1979). In contrast, Burton et al. (1999) find insignificantly positive abnormal returns for announcements of immediate and non-immediate cash-generating projects, and significantly positive abnormal returns for joint venture projects. They conclude that the market reaction to individual firm project announcements is consistent with a rational expectations explanation whereby the market anticipates the capital expenditure news of large, listed companies.

Taken together, the literature reviewed above has important, but as yet untested implications for firms undertaking investment projects with stringent environmental requirements. This study proposes that the time and costs to obtain regulatory approval for a new project may confer an economic benefit on permit holders. We test this proposition by examining the influence of expected resource consent compliance costs on the shareholder wealth impact of capital expenditure announcements.

The first hypothesis involves testing the validity of a constructed indicator of expected resource consent compliance costs. As the time to obtain consent approval increases, total compliance costs can be expected to increase (Office of the Associate Minister for the Environment, 2004). Hence the expected time to gain resource consent approval can be employed by investors as a gauge of future resource consent compliance costs. If our constructed measure of the expected time to obtain consent approval is an appropriate indicator of future regulatory delay, then according to hypothesis *H1,* a positive relationship between the two is expected.

H1: The expected time for a project to gain resource consent approval is positively related to its actual time to consent or abandon.

Following Woolridge and Snow (1990), we present shareholder value maximisation hypotheses that suggest that the expected net present value of a given project should be reflected in abnormal returns arising on the project announcement day. If a firm is able to undertake a strategic investment that enables it to gain a competitive advantage, then the expected project NPV would be greater than zero. Hence, the shareholder value maximisation hypothesis (*H2*) suggests that investors will react positively to announcements of new capital expenditures. Furthermore we posit that if firms' investments in projects with higher resource consent compliance costs allow them to create sustainable competitive advantages, then high-consent cost projects will have larger positive abnormal returns (*H2a*).

H2: Shareholder wealth maximisation hypothesis: The stock market reaction to capital expenditure project announcements is positive.

H2a: The event-window abnormal returns are greater for projects with higher expected resource consent compliance costs.

3. Data and methodology

3.1 Data, data sources and sample selection

To test the hypotheses, announcements of new capital expenditure projects undertaken by firms listed on the NZX are collected from the NZX Company Research and IRG Deep Archive databases between January 1991 and August 2007.[5] Projects are considered for inclusion in the sample if a resource consent to undertake or operate a project is either required or already possessed. Capital expenditure projects are defined as the acquisition or construction of new plant and equipment, and the upgrade of existing tangible capital assets.[6]

From the search of stock exchange announcements, 128 capital expenditure projects are identified. For each project, the initial announcement date is chosen as the earlier of the announcement of the project or the announcement of the resource consent plans.[7] To ensure that the date of the initial announcement is correctly identified, news of each identified project is also searched via the Newztext Plus database, which includes full text

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[5] The cut-off date is set to avoid the possible influence on results of a September 2007 government announcement of the planned introduction of an emissions trading scheme.

[6] Keyword search terms include 'purchase', 'develop', 'development', 'acquire' and 'acquisition' together with 'consent', 'notify', 'non-notified', and variations thereof. Following previous studies, the definition of new capital expenditures includes those projects undertaken through joint venture arrangements, but excludes marketable securities acquired through mergers and takeovers (Burton, et al., 1999, Del Brio, et al., 2003).

[7] Stock exchange announcements after market close are deemed to arise on the next working day.

coverage of NZ newspaper, newswire and magazine reports.[8] Similar to Burton et al. (1999), the earliest of the NZX announcement or media reporting date is designated as the event day, focusing upon a two-day (0,+1) event window.

To be included in the sample, each announcement is required to meet the following restrictions. First, it must be an initial announcement of the proposal or plan to undertake a capital expenditure and/or pursue resource consent approval for which the initiation date can be clearly identified. Second, no confounding events must occur within plus or minus two days of the announcement (-2,+2). Third, announcing firms' stock must have traded around the time of the announcement (-1,+1). The application of these screening criteria eliminates 73 projects from consideration, resulting in a sample of 55 non-contaminated announcements by 27 listed companies from August 1992 to July 2007. As presented in Panel A of Table 1, the greatest number of project announcements in any year is 6, reflecting high economic growth in both 2000 and 2004.

Insert Table 1 about here

In order to test hypothesis *H2a* concerning the valuation impact of resource consent compliance costs, a composite variable is constructed as explained in section 3.2 to estimate regulatory delay as the expected time (in months) to gain resource consent approval (*ETC*). Panel A of Table 1 classifies the projects annually as having a short or long *ETC*, relative to the median *ETC* of 11.15 months for the overall sample. Few *ETC* estimates are able to be made for projects early in the sample period due to a lack of relevant public information at that time. Consequently, *ETC* estimates are available for 46 out of the 55 projects, over a period from 1993 to 2007.

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[8] As Newztext Plus has limited coverage prior to 1995, the Factiva media database is also used to identify pre-1995 initial announcements.

Panel A of Table 1 also reports that 37 (67%) of the 55 project announcements make explicit mention of the related resource consent or consenting process. Virtually all major capital expansion and development projects require resource consent approval under the RMA, so for the remaining 18 announcements which failed to mention the resource consent at the time of the project initiation, in the minds of investors, the need for consent approval would still be implicit.[9] The industry affiliations of the companies represented in the sample, grouped according to the level 2 Datastream Global Industry Classifications, are presented in Panel B of Table 1. Announcements from the utilities industry make up the greatest portion of the sample at 34%, followed by industrials at 20% and financial services (property investment) at 18%. Overall, the sample companies reflect a wide range of capital-intensive industries.

To conduct the event study, we construct an equal-weighted stock market index, which avoids the problems caused by the dominance of a few large companies in the NZSE40 value-weighted equity index. The index is constructed using the Datastream live and delisted stock return indices series from 1991 to 2007 and, for a few stocks missing price and/or volume data, from the New Zealand Exchange.[10] The announcement sample daily stock returns were then captured from the stock market index dataset.

For the cross-sectional regression analysis, annual firm financial data and number of shares on issue are obtained from NZX Company Research. Market value of equity, market value of assets, trading volume and industry classifications are sourced from Datastream. Share ownership information is from NZX Company Research and company annual reports.

[9] For all of the latter projects, consenting information was disseminated through the media and/or the stock exchange at some stage prior to the commencement of operations.

[10] The index excludes 28 stocks that on average failed to trade on at least 40% of trading days to avoid estimation problems caused by thin trading (Scholes and Williams, 1977).

Additional project-specific information including the capital investment size, joint venture arrangements and resource consent details are obtained from the media or stock exchange announcements.

3.2 Measurement of expected time to gain resource consent approval

To test the shareholder wealth maximisation hypothesis in *H2a*, a measure is needed of expected resource consent compliance costs at the time of each project initiation. In a 2004 Cabinet briefing paper (Office of the Associate Minister for the Environment, 2004), the Ministry asserts that increases in compliance costs occur when consent-processing time limits are exceeded, public consultation is required, decisions are appealed to the Courts, or the duration of the consent is reduced. The paper indicates that for large, complex projects, the consenting process can impose delays that increase project-related holding and opportunity costs, and states that " the time an application takes to be granted is a useful indicator of compliance costs under the RMA" (Office of the Associate Minister for the Environment, 2004, p. 6).

To consider this point further, Table 2 outlines in chronological order the various sources of resource consent compliance costs and classifies them as time varying, time invariant or mixed. Timing-varying consent costs increase as consent-processing time increases and include consultants' fees, opportunity costs of employees' time, council hearing-related fees, costs of gathering additional information, consultation costs, delay-related holding costs and opportunity costs of poor equipment utilisation. Time-invariant consent costs do not increase with time, being the opportunity costs of Court-ordered changes in consent conditions and mitigation action costs. Finally, council application-related fees, legal costs of appeals, and environmental compliance monitoring costs are mixed costs as they contain both time-varying and time-invariant elements. Ministry for the Environment research suggests that costs arising from uncertainties over the timing of consent approval can be great, while administrative processing charges are relatively small (Quality Planning, Undated).

Insert Table 2 about here

The foregoing evidence suggests that the time-varying costs associated with the time required to gain resource consent approval are particularly material, and supports the view that the time to gain resource consent approval can be used as an indicator of resource consent compliance costs. Accordingly, we construct a composite variable, *ETC*, to estimate the expected time (in months) to gain consent approval based upon public information that investors could reasonably be expected to use at the project announcement date. To do this, we compile three forecast measures, summarised in Table 3, and detailed below.[11] Except as noted, the Table 3 statistics exclude 8 capital acquisitions for which the approved resource consent was purchased together with the capital asset (*ETC*=0).

Insert Table 3 about here

The first measure is the management forecast of *ETC*. Eleven of the 55 project announcements report a management forecast of *ETC* greater than zero, with the mean (median) being 11.55 (7.00) months. A second, time-varying measure of the historical firm-level median time to consent or abandon (*TCA*) is constructed based upon the consent processing time elapsed for past projects identified in the keyword search. This results in historical firm-level *TCA* measures for 22 out of the 55 projects in the sample, with a mean (median) of 23.08 (12.30) months. The third measure, the management forecast of the expected time for the project to become operational, is available for 22 of the 55 projects. While this measure overstates the expected time to obtain consent approval, it incorporates

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[11] A fourth, industry-level forecast is also compiled, but has no predictive power of project-level time to gain consent approval, possibly due to the considerable variation of firm-level consenting times within each industry and small industry sample sizes.

a component that allows for the resource consent processing period. Table 3 shows that the mean (median) forecast time to commence operations is 25.23 (24.00) months.

To evaluate which of the forecast measures to incorporate into the *ETC* variable for the 55 sample projects, information is needed regarding the relationship between the forecast versus actual time to consent/abandon at the project level. Table 3 reveals that the mean (median) historical project-level *TCA* is 21.45 (13.81) months. Table 3 also shows that the forecast measure most highly correlated with the historical project-level *TCA* is the management forecast *ETC* (ρ=0.8569, n=9), followed by the historical firm-level *TCA* (ρ=0.2787, n=22), and lastly the management forecast months to operate (ρ=0.1334, n=18).[12]

To emulate investors' estimates of the *ETC*, the composite variable, *ETC*, is constructed as follows for any given project in the sample:

1. If the project is already consented at the time of the project initiation announcement, then *ETC*=0. Eight projects fall into this category.

2. If *ETC* is not equal to zero, and if it is disclosed at the time of a project initiation, use the firm management forecast of *ETC*. Eleven estimates resulted from this step.

3. If an estimate of *ETC* is not available from the application of steps 1 or 2, then use the time-varying estimate of historical firm-level median *TCA* for past capital expenditure projects. Seventeen estimates were gained from this measure.

[12] No change in the rankings results when the correlations are repeated using logarithmic transformations of the variables.

4. If an estimate of *ETC* is not available from the application of steps 1, 2 or 3, then use the firm management forecast of the expected time for the project to become operational. Another 10 estimates resulted from this step.

The *ETC* variable provides an estimate of the expected time in months to consent/abandon for 46 of the 55 total sample projects. Table 3 shows the mean (median) for the *ETC* variable is 18.19 (11.15) months. The relevant data for the non-zero *ETC* observations is also presented.[13] The *ETC* and historical project-level *TCA* variables are highly non-normal, so we take the natural log of the 41 non-zero observations of the historical project-level *TCA* (*LNTCA*) and construct a dummy variable for the *ETC* measure. The sample is partitioned at the median *ETC* of 11.2 months such that *ETCDUM* is equal to 1 if the project *ETC* is above the sample median *ETC*, and zero otherwise. The correlation coefficient between *ETCDUM* and *LNTCA* is $\rho=0.4148$. If the *ETCDUM* variable is an appropriate indicator of the expected time for a given project to gain resource consent approval, then as postulated in hypothesis *H1*, it is expected that there will be a significantly positive relationship between *ETCDUM* and *LNTCA*. To test this relationship, cross-sectional regression analysis is performed using the following model.

$$LNTCA_i = \beta_0 + \beta_1 ETCDUM_i + \beta_2 LNMVA_i + \beta_3 INV/BVA_i + \beta_4 REFORMDUM_i + e_i \qquad \textbf{(1)}$$

where $LNTCA_i$ and $ETCDUM_i$ are as described above and the control variables are defined below.

Research by the Ministry for the Environment suggests that resource consent compliance costs are significant for large and complex projects, although the point is also made that the

[13] The sample size for the correlation coefficient between non-zero *ETC* and historical project-level *TCA* is small as there are only 33 observations which have both *ETC* greater than zero and historical project-level *TCA* information.

“cost of approvals is not proportional to the business size” (Office of the Associate Minister for the Environment, 2004, p. 7). Size is controlled in two ways. First, following Chen and Ho (1997) the natural log of the market value of assets ($LNMVA_i$) is used as a proxy for firm size. Second, the relative size of the project is considered by calculating, where available, the dollar value of the investment divided by the book value of firm assets (INV/BVA_i) (Chen, 2006). Given that the Ministry research fails to reach a conclusion with regard to the relationship between compliance costs and firm or project size, we make no prediction regarding the direction of the relationship between $LNTCA_i$ and the two size variables.

Major RMA legislative reforms in 2003 may have achieved their aim of reducing the costs and delays associated with the consenting process. Alternatively, if the public has become more involved in the consultation process over time, then it is possible that any efficiency gains achieved through the legislation have been offset through increased consultation costs. To test for the impact of the legislative reforms on $LNTCA_i$ this study uses a dummy variable $REFORMDUM_i$ that takes the value of one for announcements in the post-reform period (after December 2002) and zero otherwise. No prediction is made regarding the sign of the $REFORMDUM_i$ coefficient.

3.3 Event study methodology

For the tests of the remaining hypotheses, event study methodology is used to evaluate abnormal returns around capital expenditure announcements. Using a sample period of 121 days (-110,+10), abnormal returns are calculated as the difference between expected and observed market model returns over the event window.[14] The Scholes-Williams (1977) beta estimator is used to avoid the understatement of beta coefficients in the presence of infrequent trading. The estimated betas range from -0.25 to 2.29, with the median systematic

[14] The market model is widely used in event studies and is appropriately specified with sample sizes as small as 50 (Brown and Warner, 1985, Corrado and Truong, 2008).

risk of the long *ETCDUM* group being over double that of the short *ETCDUM* group. This suggests that companies undertaking projects subject to longer regulatory delays are likely to be riskier than companies pursuing capital investments with little expected regulatory delay.[15]

Unreported analysis indicates that the sample distributions of security returns and abnormal returns violate the normality assumptions of parametric tests. Consequently results are reported using the non-parametric variance-adjusted rank test (T_{CZ}) that is free from distributional assumptions and well specified in the presence of nonnormality, thin trading and event-induced variance increases (Corrado and Zivney, 1992). For robustness, we also report the results using the Patell (1976) standardised abnormal return test, the Boehmer et al. (1991) standardised cross-sectional test, and the Corrado (1989) (non-variance-adjusted) rank test. The two-sample Wilcoxon Z-test is used to test for differences between the two *ETC* subsample standardised abnormal returns.

In order to gather further evidence regarding hypothesis *H2a* with respect to the impact of expected resource consent compliance costs on project announcement abnormal returns, cross-sectional regression tests are conducted. The two-day (0,+1) and three-day (0,+2) abnormal returns for each event *i*, CAR_i, are regressed against the *ETCDUM* indicator variable and several control variables using the following ordinary least squares regression model:

$$CAR_i = \beta_0 + \beta_1 ETCDUM_i + \beta_2 LNMVA_i + \beta_3 INV/BVA_i + \beta_4 RCDUM_i + \beta_5 DISCLOSDUM_i + \beta_6 OWNCON_i + e_i \quad \textbf{(2)}$$

[15] The types of projects announced tend to be “typical” for each firm and therefore are likely to be of average risk relative to the existing firm assets.

where CAR_i *and* $ETCDUM_i$ are described above, and the control variables are defined below.

The differential information hypothesis suggests that the relatively greater attention of the media to large firms lessens the surprise element of large firm announcements, such that there is a negative relationship between firm size and event-period abnormal returns (Atiase, 1985). Firm size is measured as the natural log of the market value of firm assets ($LNMVA_i$) and a negative relationship is expected with the stock market reaction (Chen and Ho, 1997). To control for relative project size we use the dollar value of the investment divided by the book value of assets (INV/BVA_i), and predict a positive relationship with event-window abnormal returns (Chen, 2006).

All the projects in the sample require (or already hold) resource consent approval in order to go ahead, however, only two-thirds of the announcements make explicit mention of the consent. Investors can be expected to be sufficiently aware of NZ laws to understand that resource consent approval is required for major projects, however, it is possible that the additional information transmitted in announcements that discuss resource consents is valued by the market. Accordingly, the model includes a dummy variable, $RCDUM_i$ equal to one when resource consent information is explicitly disclosed in the project announcement, and zero otherwise. If the resource consent disclosure is informative, then a positive relationship is expected between $RCDUM_i$ and announcement abnormal returns.

For 24 of the 55 sample announcements, the possibility of the project is conjectured (usually through the media) prior to the formal project announcement. The prior dissemination of information may reduce informational frictions if the surprise element of a subsequent announcement is diminished, thereby reducing the follow-on stock price reaction (Palepu, 1986). Alternatively, the positive feedback theory implies that speculative news and stock price increases attract investor attention, thereby generating positive investor sentiment which in turn drives further stock demand and further price increases (Shiller, 2003). Accordingly, to test the influence of investors' anticipation of the project initiation announcements, this study uses a dummy variable *DISCLOSDUM*$_i$ which equals one for a first disclosure, and zero otherwise. No prediction for the direction of the coefficient on this variable is proposed due to the contrasting possibilities suggested by the literature.

Corporate governance literature suggests that the presence of large shareholders may affect agency relationships and firm values (Shleifer and Vishny, 1986). For example, firm financial performance may be positively affected if minority block ownership reduces management/shareholder agency conflicts, but may be negatively impacted at very high levels of ownership concentration if managers become entrenched (Thomsen and Pedersen, 2000). The ownership of NZ listed companies tends to be highly concentrated, (Gunasekarage and Reed, 2008, Hossain, et al., 2001), so in this study the proportion of equity held by the ten largest shareholders prior to the announcement day, *OWNCON*$_i$, is used to control for the possible impact of ownership concentration on announcement abnormal returns. Given that the mean (median) sample ownership concentration is extremely high at 73.8% (76.9%), a negative coefficient is expected.

4. Empirical results

4.1 Measurement of expected time to gain resource consent approval

The first test involves the validation of a constructed measure of expected consent compliance costs (*ETCDUM*). Table 4 presents the results of the OLS regression analyses to test hypothesis *H1* that the expected time for a project to gain resource consent approval is positively related to its actual time to consent or abandon (*TCA*). To overcome the problem of heteroskedasticity of disturbance terms, the White (1980) error correction method is employed. In all models, the coefficient on the *ETCDUM* variable is positive and strongly statistically significant at the 1% level, indicating that it is a valid predictor of the actual time to gain consent approval or abandon a project. The coefficients on the control variables of firm size (*LNMVA*), relative project size (*INV/BVA*), and reform (*REFORMDUM*) are statistically insignificant in all models. Untabulated analysis indicates that the events in the regression sample are fairly well distributed between the short and long *ETCDUM* categories for each industry, leading us to conclude that it is the expected time to consent measure, and not the industry affiliation, that is predicting the actual time to obtain consent approval. These results indicate support for hypothesis *H1*, and give some validation of the *ETCDUM* variable as a predictor of a project's actual time to consent or abandon. Accordingly, in subsequent regressions of project announcement *CAR*, we use the constructed *ETCDUM* variable as an indicator of expected resource consent compliance costs.

Insert Table 4 about here

4.2 The market reaction to capital expenditure announcements

Table 5 reports the event study results of the analysis of abnormal returns. The daily results around the event day for the entire sample reported in Panel A reveal weak and inconsistent evidence of abnormal returns on day 1. The cumulative abnormal returns are also given for various event windows. The two-day (0,+1) and three-day (0,+2) mean (median) *CAR* are 0.73% (0.41%) and 0.86% (0.24%), respectively (significant at the 1% and 5% levels for most of the reported test statistics). These results provide moderate evidence that capital expenditure project announcements are associated with positive valuation effects. The magnitude of the two-day results is broadly similar to those observed in comparable studies of capital investment announcements. For example, two-day mean *CAR* for similar capital investment announcements are found to be 0.33% in the US (Woolridge and Snow, 1990), 0.86% in Singapore (Chen and Ho, 1997), 0.39% in the UK (Burton, et al., 1999) and 0.30% in Korea (Kim, et al., 2005).

Insert Table 5 about here

If, counter to our predictions, the market expects new projects to have net negative valuation implications when a firm incurs high resource consent compliance costs, then the inclusion in the sample of new projects for which resource consents have already been granted (*ETC*=0) is likely to bias the results upward, as the compliance costs associated with granting the resource consents have already been incurred. Panel B of Table 5 reveals that removal from the sample of the 8 project initiations with *ETC*=0 has no material impact on the Panel A results. Furthermore, untabulated analyses find the results in Table 5 are materially unchanged when the analyses are repeated using market-adjusted returns, a more liquid market index, and a longer estimation period.

The evidence presented above provides moderate support for the shareholder wealth maximisation hypothesis *H2* that the stock market reaction to capital expenditure project announcements is positive. The findings suggest that new projects are valued positively by the market.

4.3 The influence of expected resource consent compliance costs on the market reaction to capital expenditure announcements

In order to test hypothesis *H2a* with respect to the influence of expected resource consent compliance costs on market reactions to project announcements, the sample is divided by the median *ETC* in Table 6. Projects with a short *ETC* (Panel A) have insignificant mean (median) *CAR* (0,+1) of -0.14% (0.11%) and *CAR* (0,+2) of -0.06% (-0.05%). In contrast, the two-day (0,+1) and three-day (0,+2) mean (median) *CAR* for projects with long *ETC* (Panel B) are 1.39% (0.74%) and 1.33% (1.05%), respectively, being statistically significant at the 5% level for the Corrado-Zivney rank test, and at the 5% or 1% levels for the robustness tests.

Insert Table 6 about here

Figure 1 plots the full sample and *ETC* subsample *CAR* for 20 days around the project announcements. The *CAR* for the long *ETC* subsample have an overall upward trend commencing several days before the project announcement, peaking on day +1, and then gradually trending downward after that. By day +10, the mean (median) 21-day *CAR* remain positive at 1.65% (2.3%). In contrast, the *CAR* for the short *ETC* group trend downwards in advance of the announcement from day -6 until day 0, from which point forward no particular trend is discernable. In both subsamples, much of market reaction precedes the announcement, suggesting that some of the announcements are anticipated by some investors.

Insert Figure 1 about here

Table 7 reports t-tests and Wilcoxon rank sum tests for the differences between the short and long *ETC* group standardised abnormal returns. The test statistics for the differences in the *CAR* are consistently negative, indicating that the *CAR* for the short *ETC* group are less than those for the long *ETC* group. The t-statistics are statistically significant at the 5% level for the (0,+1) and (0,+2) windows. However the Wilcoxon rank sum test for equality of medians indicates that differences are statistically significant (at the 5% level) only for the (0,+2) window.

Insert Table 7 about here

To provide further evidence regarding hypothesis *H2a*, Table 8 presents cross-sectional regression analyses of project announcement *CAR* for the (0,+1) and (0,+2) windows. The t-statistics are calculated using White (1980) heteroskedasticity-consistent standard errors. In Models 1 and 2, *ETCDUM* is the sole explanatory variable, and the coefficients are positive but not significant at conventional levels. They remain insignificant when the explanatory variables for firm size (*LNMVA*) and relative project size (*INV/BVA*) are added in Models 3 and 4. The negative coefficient (1% level) for *LNMVA* is consistent with similar studies suggesting that small firms experience greater information asymmetry (Chen, 2006, Chen and Ho, 1997). The sample size for Models 3 and 4 is relatively small as the project size is disclosed for only 32 out of the 46 announcements for which an *ETCDUM* estimate is available. Given the data limitations and the insignificant t-statistics associated with project size in the models, it is dropped from consideration in subsequent models.

Insert Table 8 about here

Models 5 and 6 add three further control variables, *RCDUM, DISCLOSDUM* and *OWNCON*. In these models, the coefficients for *ETCDUM* are positive and significant at the 1% and 5% levels in the (0,+1) and (0,+2) windows, respectively.[16] The t-statistic for *RCDUM* is positive and statistically significant at the 5% level in the *CAR* (0,+1) model, suggesting that resource consent information is positively valued, possibly due to uncertainty reduction. If the announcement is a first disclosure then according to the significantly negative coefficients (5% level) on *DISCLOSDUM*, then the market reaction is negative. Conversely, anticipated announcements generate more positive reactions than non-anticipated (first) announcements, consistent with the positive feedback theory (Shiller, 2003) which suggests that speculative news attracts investors' attention generating positive sentiment which drives further share price increases. For *OWNCON*, the coefficient estimate is significantly negative (5% level) as predicted, but only in the *CAR* (0.+1) model. Consistent with prior research (e.g. Hossain, et al., 2001, Thomsen and Pedersen, 2000), this suggests that extremely high block ownership may negatively affect firm performance.

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[16] These analyses implicitly assume a linear relationship between the expected time to obtain consent and the investment value to shareholders. Alternate specifications of equation 2, including the use of *ETC* and ETC^2 in place of *ETCDUM,* reveal no evidence of non-linearity.

One possible explanation for the finding of a higher event window *CAR* for long time-to-consent projects is that investors assign different probabilities to the likelihood of a firm announcement for the two *ETCDUM* groups. Pre-announcement conjecture concerning possible projects could lead investors to assign a higher probability of an announcement for short time-to-consent projects relative to long time-to-consent projects. In such a case, it is possible that the market reaction for the short *ETC* group arises at the time of the pre-announcement information dissemination, and there is no abnormal return in response to the subsequent project initiation announcement. We test for this possibility in several ways (untabulated). Firstly, we add to cross-sectional regression equation (2) an interaction term between the *ETCDUM* and *DISCLOSDUM* variables. The coefficient is statistically no different from zero, suggesting that there is no differential effect of anticipation between the two *ETCDUM* groups due to pre-announcement disclosures. Secondly, given that the subsample sizes are small, we perform Wilcoxon rank sum exact tests to check for significant differences between the event window *CAR* for anticipated and non-anticipated announcements for the short *ETC* group. No significant difference is found, suggesting than anticipation does not affect *CAR* differently when the expected time to consent is short. Nevertheless, some media conjecture preceded the announcement for three out of eight projects in the short *ETC* group for which approved resource consents were purchased together with the capital assets. For these three projects, market participants were aware that the projects were already operating with resource consents in place. If investors factor the value of this information into the stock price prior to the announcement, then the short *ETC* group event study findings in Panel A of Table 6 could be biased downwards. The deletion of these three announcements from the short *ETC* group results in no material change to the reported results, suggesting that the findings are not driven by anticipated projects with purchased consents.

Another potential factor that could influence the interpretation of results is the relative probability of gaining resource consent approval for each project. If for example, a lower probability of gaining approval enhances firms' abilities to develop competitive advantages, then any difference in the average expected probability of approval between the two *ETCDUM* groups could affect the interpretation of the results. In each *ETCDUM* group in this study, there is one project that was abandoned without consent approval and one project for which approval could not be confirmed. Accordingly, there is no evidence to suggest that the probability of consent approval differs between the groups. In the larger group of 128 capital expenditure projects initially identified, the probability of gaining consent approval lies between 89.8% and 96.9%.[17] Given that actual consent approval is relatively high overall and is identical between the *ETCDUM* groups, the probability of consent approval appears unlikely to have materially affected the results.

In further (untabulated) analyses, we test the robustness of the cross-sectional regression results. We use dummy variables to test for possible effects from 2003 RMA reforms, the presence of joint venture arrangements (Burton, et al., 1999) and energy generation projects. We also check for the possible influence of liquidity effects due to thin-trading (Anderson, et al., 2006), as well as growth opportunities and financial leverage on our regression results (Chen and Ho, 1997). The analysis is also repeated deleting a potentially important outlier. Overall, the results are robust to these modifications.

The above findings are consistent with the shareholder wealth maximisation hypothesis *H2* and the corollary predictions in *H2a*, that the event-window abnormal returns are greater for projects with higher expected resource consent compliance costs. The regression results indicate that the average values of capital expenditure projects expected to experience long

[17] Four projects are identified as being abandoned without consent approval, no record of a consent decision can be found for another five projects, and another four projects were sold with no record found of a consent decision.

consenting processes exceed those with short consenting processes by about 1.51% to 2.00% (using *CAR* (0,+1) as the dependent variable). Applying these percentages to the market value of each announcer's equity at the fiscal year-end prior to the announcement, the estimated average announcement net benefit expressed in 2007 dollars is $18.0 to $23.8 million (relative to the average market value of equity of $1,189.3 million).[18] These results suggest that on average, the marginal expected benefits of undertaking environmentally-sensitive projects that are expected to experience relatively lengthy consent processing delays are substantially greater than the marginal expected compliance costs.

The above findings also suggest that the stock market reaction to project announcements for which environmental regulatory delays are expected to be short, is statistically no different from zero. For these projects, potential competitive advantages may be diminished if short resource consent approval times lessen firms' opportunities to benefit from early mover advantages, reputational benefits, or highly developed environmental management systems. An alternative interpretation is that lower compliance costs result in fewer barriers to impede industry competitors and new entrants, hence competition is greater and the opportunity to earn economic profits is diminished.

An important caveat should be acknowledged when interpreting the findings. The total sample size is small (n=55), subsample analyses are conducted with even smaller groups, and only limited data on relative project size is available. Although the non-parametric tests used for the analyses are robust to small sample sizes, the number of explanatory variables able to be included simultaneously in the cross-sectional regression models is limited. Control variables are added to the models consecutively to mitigate the small sample

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[18] In 1991 dollars, this translates to a net benefit of $13.0 to $17.2 million (relative to the average market value of equity of $860.8 million).

problem, but it is possible that the results could be different if all explanatory variables could be incorporated simultaneously or if relative project size data was more readily available.

5. Conclusion

This study investigates the influence of environmental regulatory delay in explaining the capital market impact of NZ capital expenditures. Consistent with the shareholder wealth maximisation hypothesis, we find that overall, the stock market positively values news of capital expenditure projects. However, the positive valuation is driven by those project announcements for which the expected time to obtain resource consent approval is long, which is consistent with the view that environmental regulations may enhance the wealth of regulated firms (e.g. Maloney and McCormick, 1982, Porter and van der Linde, 1995). We suggest that the time delays for consent approval and high level of compliance costs incurred for environmentally-sensitive projects may allow firms to develop specialised capabilities and/or to deter industry competitors and new entrants, thereby increasing expected project NPVs.

Our analysis also offers insights into the nature of nonfinancial information used by investors in their stock pricing decisions. The results suggest that investors use consent status (i.e. already approved or not), management forecasts of delays and information about firm-level regulatory delays from past projects to estimate future regulatory compliance costs associated with proposed projects. The implication that investors attach importance to management forecasts of delays is consistent with prior evidence that analyst reports and EPS forecast revisions are frequently triggered by forward-looking management announcements (Kerl, et al., 2012). Given that publicly-available analyst recommendations are a major driver of stock price changes (Ryan and Taffler, 2004), the apparent (indirect) association between management forecasts of delays and firm value appears credible. The findings that compliance cost indicators, nonfinancial consent disclosures and project-related financial information result in significant abnormal returns around the days of capital expenditure announcements contribute to research evidence that financial and nonfinancial information communicated independently from financial statements is relevant to the valuation of firms.

The findings in this paper have important implications for regulators. If protracted environmental regulatory processes assist wealth creation through improved project decision-making and enhanced environmental risk management systems, then the benefits of environmental protection policies may be shared widely across stakeholders. Furthermore, the opportunities for firms to earn economic rents may encourage capital investment. Nevertheless, an alternate explanation for our findings is that regulatory delays allow early movers the opportunity to erect barriers to entry, thereby conferring a relative advantage on incumbent firms. Consequently, environmental regulations may have the troubling consequences of decreasing market competition through redistribution of industry wealth and increases in industry concentrations (Helland and Matsuno, 2003). This implies that if legislators are able to reduce environmental regulatory delays associated with capital expenditure approval processes, then the opportunity for firms to earn economic rents may be diminished.

Our results relate to the small, NZ economy, so an international comparative study of the impact of consent/permit processes across different regulatory jurisdictions would provide a useful next research step. Much of the firm learning and early mover benefit research originates from the US (e.g. Dean and Brown, 1995, King and Lenox, 2002), implying that the results of this study may well have wider applicability. Further research could examine the effect of environmental regulatory delay on new firm entries, firm exits and relative firm competitiveness within an industry (Millimet, et al., 2009). Another potential research direction could test Hart's (1995) natural resource-based theory of the development of valuable internal capabilities by interviewing business managers to gain insights into potential sources of strategic advantages gained through firm management of environmental regulatory processes.

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Figure 1 Full sample and ETC subsample cumulative abnormal returns around project announcements

The figure below compares the cumulative abnormal returns of short, long and all expected time-to-consent (ETC) project announcements from event day -10 to +10.

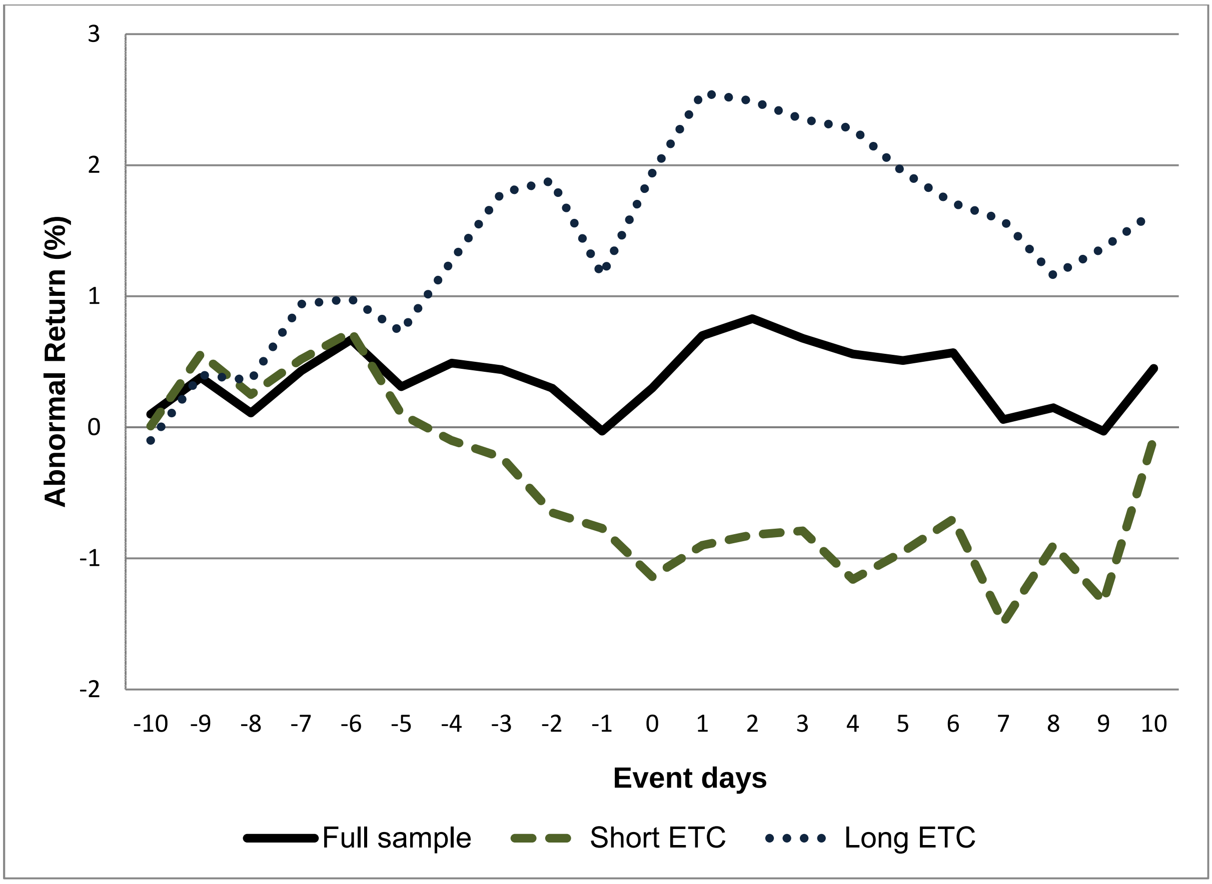

Table 1 Project sample analysis

The sample distributions of listed-company project announcements from 1992 to 2007 are summarised by year in Panel A and by industry in Panel B. In Panel A, the sample is divided by the median ETC (where available), where ETC is the expected time for a project to gain resource consent approval at the time of the project announcement. Of the 55 sample announcements, 37 explicitly mention resource consents.

Panel A. Sample of project announcements by year

	All	Short ETC (< median)	Long ETC (> median)	Resource consent mentioned
1992	2	0	0	2
1993	3	0	1	2
1994	1	0	0	0
1995	4	2	1	3
1996	3	1	1	1
1997	5	3	2	4
1998	3	1	2	1
1999	2	1	1	2
2000	6	3	1	5
2001	1	1	0	1
2002	5	1	4	2
2003	5	3	2	3
2004	6	3	3	4
2005	4	3	1	2
2006	4	1	3	4
2007	1	0	1	1
Total	55	23	23	37

Panel B. Industry affiliations of project announcement sample

Datastream Industry Classification Level 2	Number of companies	Total announcements	
		Number	Percent
Basic materials	1	2	4
Consumer goods	1	1	2
Consumer services	3	1	6
Financial services	4	10	18
Healthcare	2	6	11
Industrials	6	11	20
Oil & gas	3	3	5
Utilities	7	19	34
Total	27	55	100

Table 2 Sources of resource consent compliance costs

The sources of resource consent compliance costs are described and categorised as time-varying, time-invariant, and mixed.

Source of compliance cost	Cost type	Explanation
Consulting fees	Time-varying	This includes payments to consultants to advise and/or provide services relating to the mandatory assessment(s) of environmental effects and the resource consent application(s).
Opportunity costs of employees' time devoted to the resource consent approval process	Time-varying	Staff resources are required throughout the resource consent process.
Council application-related fees	Mixed	The time-varying portion relates to the cost of public notices and hours required by authority planners, advisors and administrators. Within each authority, the application fee is fixed (time invariant) for each type of consent. However, several applications are needed if the geographical region relates to more than one authority.
Council hearing-related fees	Time-varying	Applicants pay per-hour charges for chairperson, councillors, consultant planners, independent commissioners, compliance officers and administrative officers. Venue hire costs increase as number of hearings increases.
Costs of gathering additional information requested by consent authority	Time-varying	Consent authorities may request further information from the applicant to support their case.
Consultation costs	Time-varying	This includes consultation with interested groups and iwi.
Legal costs of appeals	Mixed	Includes lawyers' and expert witness fees, and contingent costs of court-awarded costs to submitters.
Delay-related holding and opportunity costs	Time-varying	Includes deferral of project revenue and poor utilisation of labour and expensive equipment.
Opportunity costs of changes in consent conditions	Time invariant	Appeals may result in changes to the conditions of the consents.
Mitigation action costs	Time invariant	To gain consent approval, an applicant may agree to conditions which seek to mitigate adverse impacts. While the consent conditions are known at the time of approval, the actual costs may be incurred subsequent to approval.
Monitoring costs	Mixed	To gain consent approval, an applicant may agree to conditions that require them to incur ongoing environmental compliance monitoring costs.

Sources: Office of the Associate Minister for the Environment (2004), Quality Planning (Undated),Sheppard (1998)

Table 3 Summary statistics and correlations - Expected time to consent (ETC) components and historical time to consent/abandon (TCA)

Table 3 presents summary statistics for the components considered for construction of a composite variable, ETC, to estimate the expected time (in months) to gain consent approval based upon public information at the project announcement date. Component variables reported are the management forecast of ETC, historical firm-level TCA (time to consent or abandon) and management forecast of months for the project to become operational. Also reported are summary statistics for the historical (actual) project-level TCA, the composite ETC variable, and their transformed equivalents. LNTCA (historical project-level) is the natural log of the historical project-level TCA, while ETCDUM is a dummy variable divided at the median of the ETC variable. The ETC variable is constructed from three component variables that are positively correlated with the historical project-level TCA. The Pearson correlation coefficient between each component (forecast) variable with the historical project-level TCA is reported, where a and b denote statistical significance at the 1% and 5% levels, respectively.

Variable (months)	n	Mean	Median	Std. Dev.	Min.	Max.	Correlation (ρ) with historical project-level TCA
Panel A: Non-transformed data							
Mgt forecast of ETC	11	11.55	7.00	11.70	2.00	42.00	0.8569[a] (9)
Historical firm-level TCA	22	23.08	12.30	26.65	5.29	86.10	0.2787 (22)
Mgt forecast months to operate	22	25.23	24.00	20.86	3.00	78.00	0.1334 (18)
Historical project-level TCA	41	21.45	13.81	22.71	1.64	111.78	
ETC variable (including ETC=0)	46	18.19	11.15	22.97	0.00	86.10	
ETC variable (excluding ETC=0)	38	22.01	12.00	23.56	2.00	86.10	0.1648 (33)
Panel B: Transformed variables							
LNTCA (historical project-level)	41	2.65	2.63	0.94	0.50	4.72	
ETCDUM variable (including ETC=0)	46	0.50	0.50	0.51	0.00	1.00	
ETCDUM variable (excluding ETC=0)	38	0.61	1.00	0.50	0.00	1.00	0.4148[b] (33)

Table 4 Cross-sectional regression analyses of LNTCA

$$LNTCA_i = \beta_0 + \beta_1 ETCDUM_i + \beta_2 LNMVA_i + \beta_3 INV/BVA_i + \beta_4 REFORMDUM_i + e_i \quad \textbf{(1)}$$

The table below reports the cross-sectional regression analyses of the log of the historical project-level time to consent or abandon (LNTCA) on the expected time to consent dummy variable (ETCDUM), firm size, measured as the log of market value of assets (LNMVA), project size, measured as the ratio of the investment cost of the project to the book value of firm assets (INV/BVA), and a dummy variable (REFORMDUM) denoting 1 for announcements after December 2002 RMA legislative reforms, and 0 otherwise. Coefficient estimates are presented with White (1980) heteroskedasticity-consistent p-values reported in brackets below. Tests for multicollinearity reveal no evidence of high correlation between independent variables. a and b denote statistical significance at the 1% and 5% levels, respectively.

Variable	Predicted sign	Model 1 (n=33)	Model 2 (n=33)	Model 3 (n=20)	Model 4 (n=33)	Model 5 (n=20)
Constant		2.0694 (11.82)[a]	4.4289 (2.44)[b]	2.0460 (10.30)[a]	2.1396 (11.71)[a]	3.9057 (1.90)
ETCDUM variable (excluding ETC=0)	+	0.7238 (2.72)[a]	0.7595 (2.88)[a]	0.8618 (2.78)[a]	0.7127 (2.79)[a]	1.0105 (3.16)[a]
LNMVA	n/a		-0.1176 (-1.30)			-0.0990 (-1.02)
INV/BVA	n/a			1.1072 (0.88)		0.7640 (0.47)
REFORMDUM	n/a				-0.1404 (-0.53)	0.3513 (1.10)
Adjusted R^2		0.144	0.149	0.329	0.122	0.309

Table 5 Abnormal returns and cumulative abnormal returns around project announcements

This table reports mean and median abnormal returns and mean ranked variance-adjusted standardised abnormal returns around project announcements based upon market model residuals with Scholes-Williams betas using a (-110,+10) sample period. Panel A reports results for the entire sample, and Panel B reports the sample results omitting 8 projects for which consent approval has already been granted (ETC=0). Statistical significance is evaluated using the non-parametric Corrado and Zivney (1992) variance-adjusted rank test (T_{CZ}) to test the null hypothesis that mean ranked event-day standardised abnormal returns are no different from zero. The rank test uses the standard deviation of abnormal returns over the entire sample period, so the reported standard deviation is identical for day 0 and each of the surrounding days. We also report the results using the Patell (1976) standardised abnormal return test (T_{PATELL}), the Boehmer, Musumeci and Poulsen (1991) standardised cross-sectional test (T_{BMP}), and the Corrado (1989) non-variance-adjusted rank test (T_C). a and b denote statistical significance at the 1% and 5% levels, respectively, using two-tailed tests.

	Abnormal returns			Ranked variance-adjusted standardised abnormal returns					
Event days	Mean	Median	Propn. pos. returns	Mean	Std dev	T_{CZ}	T_{PATELL}	T_{BMP}	T_C
Panel A. Entire sample (n=55)									
-2	-0.0014	-0.0003	0.49	-0.2779	0.3173	-0.88	-0.59	-0.63	-0.83
-1	-0.0033	-0.0020	0.33	-0.5424	0.3173	-1.71	-1.53	-1.58	-1.66
0	0.0033	0.0005	0.55	0.4851	0.3173	1.53	1.12	0.96	1.51
1	0.0040	0.0022	0.58	0.5882	0.3173	1.85	2.08[b]	1.99[b]	1.89
2	0.0013	0.0006	0.53	0.3838	0.3173	1.21	0.25	0.28	1.15
Event window									
(-1,0)	0.0000	-0.0016	0.44	-0.0573	0.4487	-0.13	-0.29	-0.31	-0.10
(0,+1)	0.0073	0.0041	0.58	1.0733	0.4487	2.39[b]	2.25[b]	1.97[b]	2.41[b]
(-1,+1)	0.0040	0.0011	0.56	0.5308	0.5495	0.97	0.98	0.97	1.01
(0,+2)	0.0086	0.0024	0.56	1.4570	0.5495	2.65[a]	1.96	2.14[b]	2.63[a]
(-2,+2)	0.0039	0.0079	0.56	0.6367	0.7095	0.90	0.59	0.61	0.92
Panel B. Excluding ETC=0 (n=47)									
-2	-0.0013	-0.0010	0.45	-0.2719	0.3253	-0.84	-0.63	-0.69	-0.80
-1	-0.0040	-0.0024	0.32	-0.5805	0.3253	-1.78	-1.59	-1.58	-1.82
0	0.0037	0.0005	0.55	0.4599	0.3253	1.41	1.17	0.95	1.49
1	0.0041	0.0023	0.60	0.6306	0.3253	1.94	2.26[b]	2.18[b]	1.92
2	0.0009	-0.0001	0.50	0.2640	0.3253	0.81	0.04	0.05	0.81
Event window									
(-1,0)	-0.0003	-0.0031	0.40	-0.1205	0.4600	-0.26	-0.29	-0.30	-0.23
(0,+1)	0.0078	0.0054	0.57	1.0906	0.4600	2.37[b]	2.43[b]	2.10[b]	2.41[b]
(-1,+1)	0.0038	0.0007	0.53	0.5101	0.5634	0.91	1.08	1.09	0.92
(0,+2)	0.0086	0.0007	0.53	1.3546	0.5634	2.40[b]	1.98[b]	2.08[b]	2.44[b]
(-2,+2)	0.0034	0.0066	0.53	0.5022	0.7273	0.69	0.56	0.57	0.72

Table 6 Abnormal returns and cumulative abnormal returns around project announcements by expected time to consent

This table reports short and long ETC subsample mean and median abnormal returns and mean ranked variance-adjusted standardised abnormal returns around project announcements based upon market model residuals with Scholes-Williams betas using a (-110,+10) sample period. The samples are divided according to the median ETC. Statistical significance is evaluated using the non-parametric Corrado and Zivney (1992) variance-adjusted rank test (T_{CZ}) to test the null hypothesis that mean ranked event-day standardised abnormal returns are no different from zero. The rank test uses the standard deviation of abnormal returns over the entire sample period, so the reported standard deviation is identical for day 0 and each of the surrounding days. We also report the results using the Patell (1976) standardised abnormal return test (T_{PATELL}), the Boehmer, Musumeci and Poulsen (1991) standardised cross-sectional test (T_{BMP}), and the Corrado (1989) non-variance-adjusted rank test (T_C). a and b denote statistical significance at the 1% and 5% levels, respectively, using two-tailed tests.

	Abnormal returns			Ranked variance-adjusted standardised abnormal returns					
Event days	Mean	Median	Propn. pos. returns	Mean	Std dev	T_{CZ}	T_{PATELL}	T_{BMP}	T_C
Panel A. Short ETC (n=23)									
-2	-0.0042	-0.0026	0.48	-0.2923	0.3043	-0.96	-1.06	-0.87	-1.19
-1	-0.0012	-0.0018	0.35	-0.3179	0.3043	-1.04	-0.75	-0.95	-0.52
0	-0.0037	0.0003	0.52	-0.0034	0.3043	-0.01	-1.20	-1.03	-0.03
1	0.0024	0.0021	0.61	0.3025	0.3043	0.99	0.43	0.54	0.87
2	0.0008	-0.0007	0.48	0.1453	0.3043	0.48	-0.36	-0.36	0.44
Event window									
(-1,0)	-0.0049	-0.0031	0.35	-0.3213	0.4303	-0.75	-1.38	-1.82	-0.39
(0,+1)	-0.0014	0.0011	0.52	0.2991	0.4303	0.69	-0.54	-0.59	0.59
(-1,+1)	-0.0025	-0.0006	0.48	-0.0188	0.5270	-0.04	-0.87	-1.14	0.18
(0,+2)	-0.0006	-0.0005	0.39	0.4444	0.5270	0.84	-0.67	-0.95	0.74
(-2,+2)	-0.0060	-0.0020	0.43	-0.1658	0.6804	-0.24	-1.31	-1.43	-0.20
Panel B. Long ETC (n=23)									
-2	0.0009	0.0019	0.57	-0.0137	0.2968	-0.05	0.08	0.12	-0.03
-1	-0.0072	-0.0024	0.30	-0.6102	0.2968	-2.06	-1.90	-1.86	-2.11[b]
0	0.0077	0.0005	0.52	0.3948	0.2968	1.33	1.95	1.65	1.33
1	0.0062	0.0030	0.61	0.4956	0.2968	1.67	2.33[b]	1.75	1.74
2	-0.0006	0.0011	0.57	0.2154	0.2968	0.73	0.31	0.34	0.60
Event window									
(-1,0)	0.0006	-0.0016	0.44	-0.2154	0.4197	-0.51	0.03	0.03	-0.55
(0,+1)	0.0139	0.0074	0.61	0.8905	0.4197	2.12[b]	3.06[a]	2.24[b]	2.18[b]
(-1,+1)	0.0067	0.0044	0.61	0.2803	0.5141	0.55	1.39	1.16	0.56
(0,+2)	0.0133	0.0105	0.65	1.1058	0.5141	2.15[b]	2.66[a]	2.47[b]	2.13[b]
(-2,+2)	0.0070	0.0086	0.65	0.4820	0.6636	0.73	1.24	1.29	0.69

Table 7 Comparison of standardised abnormal returns and standardised cumulative abnormal returns around project announcements by expected time to consent

This table reports t-tests and the non-parametric Wilcoxon rank sum test for significant differences in the mean and median standardised abnormal returns, respectively, between the short and long expected time to consent (ETC) groups. a and b denote statistical significance at the 1% and 5% levels, respectively.

	Subsample standardised abnormal returns				Difference	
	Short ETC n=23		Long ETC n=23		t-statistic	Wilcoxon Z
Event days	(1) Mean	(2) Median	(3) Mean	(4) Median	(1)-(3)	(2)-(4)
-2	-0.2237	-0.1169	0.0179	0.1547	-0.81	-0.92
-1	-0.1590	-0.0733	-0.4037	-0.2178	0.89	0.79
0	-0.2513	0.0126	0.4093	0.0481	-1.88	-1.32
1	0.0983	0.1139	0.5084	0.2405	-1.23	-0.90
2	-0.0746	-0.0165	0.0679	0.1001	-0.50	-0.33
Event window						
(-1,0)	-0.2903	-0.0871	-0.0002	-0.0749	-1.14	-0.90
(0,+1)	-0.1099	0.0110	0.6465	0.2877	-2.15[b]	-1.47
(-1,+1)	-0.1810	-0.0363	0.2918	0.1679	-1.55	-1.47
(0,+2)	-0.1318	-0.0114	0.5673	0.3332	-2.54[b]	-2.17[b]
(-2,+2)	-0.2720	-0.0836	0.2649	0.2880	-1.87	-1.42

Table 8 Cross-sectional regression analyses of project announcement cumulative abnormal returns, (t-statistics)

$$CAR_i = \beta_0 + \beta_1 ETCDUM_i + \beta_2 LNMVA_i + \beta_3 INV/BVA_i + \beta_4 RCDUM_i + \beta_5 DISCLOSDUM_i + \beta_6 OWNCON_i + e_i \quad (2)$$

Cross-sectional regression analyses of two-day (0,+1) and three-day (0,+2) cumulative abnormal returns (CAR) on the expected time to consent dummy (ETCDUM) and control variables are presented for the sample of project announcements. The key independent variable is the ETCDUM variable, which takes the value of 1 for long time-to-consent projects, and 0 otherwise. Control variables are described as follows. Firm size (LNMVA) is the log of market value of assets, project size (INV/BVA) is the ratio of the investment cost of the project to the book value of firm assets, RCDUM equals 1 if resource consent information is explicitly disclosed in the project announcements and 0 otherwise, DISCLOSDUM equals 1 for first disclosures, and 0 otherwise, and OWNCON is the proportion of equity held by the ten largest shareholders prior to the announcement day. Coefficient estimates are presented with White (1980) heteroskedasticity-consistent p-values reported in brackets below. Tests for multicollinearity reveal no evidence of high correlation between independent variables. a and b denote statistical significance at the 1% and 5% levels, respectively.

Variable	Predicted sign	CAR (0,+1)) Model 1 (n=46)	CAR (0,+2) Model 2 (n=46)	CAR (0,+1) Model 3 (n=32)	CAR (0,+2) Model 4 (n=32)	CAR (0,+1) Model 5 (n=46)	CAR (0,+2) Model 6 (n=46)
Constant	n/a	-0.0012 (-0.29)	-0.0003 (-0.06)	0.2455 (3.79)[a]	0.1810 (2.46)[b]	0.2969 (5.19)[a]	0.2305 (5.02)[a]
ETCDUM	+	0.0151 (1.89)	0.0136 (1.92)	0.0149 (1.97)	0.0120 (1.62)	0.0200 (2.67)[a]	0.0161 (2.48)[b]
LNMVA	-			-0.0121 (-3.83)[a]	-0.0087 (-2.49)[b]	-0.0139 (-4.84)[a]	-0.0108 (-4.76)[a]
INV/BVA	+			-0.0002 (-0.01)	0.0008 (0.02)		
RCDUM	+					0.0174 (2.07)[b]	0.0123 (1.73)
DISCLOSDUM	n/a					-0.0142 (-2.24)[b]	-0.0160 (-2.58)[b]
OWNCON	-					-0.0003 (-2.09)[b]	-0.0002 (-0.98)
Adjusted R^2		0.051	0.053	0.247	0.173	0.374	0.310