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

Carolyn Wirth, Jing Chi and Martin Young

Massey University

Carolyn Wirth C.G.Wirth@massey.ac.nz (corresponding author), Jing Chi J.Chi@massey.ac.nz, and Martin Young M.Young@massey.ac.nz, School of Economics and Finance, Massey University, Private Bag 11-222, Palmerston North +64 6 356 9099.

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Abstract

This study investigates the role of environmental regulatory delay in explaining capital market reactions to capital expenditure announcements. We use the time to gain resource consent (regulatory approval) as an indicator of expected resource consent compliance costs, and find positive valuation effects from project announcements when the expected time to consent is long. Our findings suggest that by undertaking voluntary capital expenditures with high environmental compliance costs, firms can create strategic advantages. For example, long consent times may improve firms' opportunities to develop specialised capabilities such as early mover advantages, reputational benefits, or enhanced environmental management systems. Alternatively, high environmental compliance costs may inhibit actions by industry competitors and new entrants, resulting in greater expected project NPVs.

Keywords: Capital investments, market valuation, environmental regulation, compliance costs, New Zealand

JEL Classifications: G31, Q58

1. Introduction

This study investigates the role of environmental regulatory delay in explaining capital market reactions to capital expenditure announcements by testing the shareholder wealth maximisation hypothesis. Under New Zealand environmental law, individuals and businesses must obtain resource consent approval to use resources or undertake development activities that may have potentially adverse environmental impacts. The consenting process 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), but research evidence of the actual impact on capital expenditures is sparse. The shareholder wealth maximisation hypothesis predicts that the stock market reacts favourably to capital expenditure project initiation announcements when managers undertake positive NPV projects (Miller & Modigliani, 1961). The neo-classical economics view suggests that, *ceteris paribus*, projects with lower resource consent compliance costs will be more valuable. However, following Porter's (1979) assertion that strategic capital expenditures may provide firms with competitive advantages, we hypothesise that projects with higher resource consent compliance costs may provide firms with first-mover or other sustainable advantages that make them more valuable.

To test the above hypothesis, we use the forecast time for each project to obtain consent as an indicator of expected resource consent compliance costs at the time of each project announcement. We find our forecast indicator to be positively related to the actual time taken to either obtain consent or abandon the project. Consistent with the shareholder wealth maximisation hypothesis, we document a positive stock market reaction to a sample of New Zealand project announcements from 1992 to 2007. Dividing the sample at the median expected time to consent, we find that the positive valuation effects are significant only for those project announcements for which the expected time to consent is long. This suggests that by undertaking voluntary capital expenditures with high environmental compliance costs, firms can gain a source of strategic advantage, possibly through the development of valuable firm-specific capabilities (Hart, 1995) or through the imposition of barriers to entry (e.g. Dean & Brown, 1995; Ryan, 2005). For these projects, competitive advantages may be greater if regulatory delays enhance firms' opportunities to develop specialised capabilities and resources, such as early mover advantages, reputational benefits, or sophisticated environmental management systems. Alternatively, the level of resource consent compliance costs incurred may be sufficient to inhibit actions by industry competitors and new entrants, resulting in greater expected project NPVs.

Our research contributes to the literature in several ways. Contrary to the negative attention that resource consent 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 & McCormick, 1982). 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. Compared with compliance cost surveys that only measure costs, our focus on capital expenditures announcements allows insights into net benefits at the project-level. Our study also directly addresses the criticism by Jaffee, Peterson, Portney and Stavins (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. We document a positive economic impact of regulatory delays in connection with capital expenditure decisions made by firms. Finally, these findings add to the growing research evidence on the valuation implications of environmental information.

The paper is structured as described next. Section 2 outlines New Zealand environmental law and the controversies surrounding the resource consent process. Section 3 provides a brief review of prior literature and develops the hypotheses. The data and methodologies employed are discussed in Section 4, and the empirical results are presented in Section 5. Concluding remarks are made in Section 6.

2. Institutional background

The Resource Management Act (RMA, 1991) was introduced in an effort to promote sustainable management of resources and to control potentially adverse environmental effects of activities on land, air and water. The legislation requires individuals and businesses to apply for resource consent 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. 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.

Notwithstanding a dearth of empirical research on the business impact of the RMA, the resource consent process has provoked a considerably amount of controversy. Consent processes have been criticised as inconsistent and costly because local consent authorities lacked the funding, policy guidance and expertise needed to ensure efficient administration (Ministerial Panel on Business Compliance Costs, 2001; OECD, 1996, 2007). Councils are argued to have overstepped their authority and been overly risk averse when considering resource consents (Upton, 1997). Furthermore, nuisance objectors and trade competitors are said to have unduly delayed the consultative stages of consent processes (Ernst & Young, 1997; Ministerial Panel on Business Compliance Costs, 2001). Some research evidence appears to support the view that protracted resource consent processes may have inhibited business investment from the early 1990s to 2001 (Clough, 2005; Wilkinson, 2001).

Many of the above problems have been blamed as the cause of onerous business compliance costs, but few studies have attempted to measure them. A 1997 survey by Ernst & Young (1997) of 73 businesses that had applied for resource consent found that compliance costs amounted to less than 5% of total project costs for the majority of businesses. However, in a 2002 survey of 500 firms, fewer than 40% of respondents considered that the benefits of gaining resource consents either exceeded or equalled the costs (Massey, 2003). A more recent survey reports that from 2006 to 2008 environmental compliance costs as a percentage of sales rose from 0.2% to 0.3% for the largest firms surveyed (Business New Zealand & KPMG, 2008).

Furthermore, there appear to be impediments to the disclosure of environmental compliance costs 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. Tozer and Hawkes (2001) reveal that out of 24 New Zealand companies surveyed that held resource consents, only one reported them separately in the financial statements. 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, Krishnan, & Lave, 2001). Taken together, the above discussion implies that while resource consent costs may be particularly material, businesses and their investors face a challenge to identify their financial implications.

3. Literature review and hypothesis development

The criticisms raised of excessive resource consent compliance costs and uncertainties for businesses that undertake capital expenditures are consistent with the neoclassical economics view that environmental regulations impose a net cost upon businesses. According to this view, environmental capital expenditures are likely to result in either negative returns to firms, or at least returns that are substantially lower than are possible from other investments (e.g. Palmer, Oates, & Portney, 1995; Walley & Whitehead, 1994). Consistent with this proposition, some studies find the stock market reacts negatively to environmental legislative news announcements (e.g. Blacconiere & Northcut, 1997; Shane, 1995)

In contrast, the 'Porter hypothesis' (Porter & van der Linde, 1995) suggests that firms may be able to gain net economic benefits through well-designed environmental regulations that encourage resource productivity, enhance innovation and improve competitiveness. Some of the literature supporting this view suggests that the allocation of property rights (such as 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 & Tullock, 1975; Maloney & McCormick, 1982; Mason & Polasky, 1994). An alternative body of literature has developed from Hart's (1995) 'natural-resource-based view of the firm' which holds that valuable internally-developed capabilities and resources within the firm may provide sustainable competitive advantages. Superior environmental performance has been positively associated with improved operational performance (Melnyk, Sroufe, & Calantone, 2003), profitability (Russo & Fouts, 1997) and firm value (King & Lenox, 2002). In contrast, other studies find evidence that the net benefits are not sufficient to overcome the direct and indirect environmental compliances costs (Jaggi & Freedman, 1992; King & Lenox, 2001). Similarly, while some studies show that sunk costs caused by the irreversibility of environmental capital expenditures impose barriers to new firm entry (e.g. Dean & Brown, 1995; Ryan, 2005), evidence on potential benefits is mixed. Environmental capital outlays mandated by regulatory requirements may hamper firm productivity and reduce firm value (Barbera & McConnell, 1986; Johnston, 2005). However, first movers, low polluters, and firms undertaking voluntary environmental capital expenditures may derive economic advantages from their investment activities (Clarkson, Li, & Richardson, 2004; Johnston, 2005; Nehrt, 1996) .

From the above literature we can infer that firms may gain first-mover advantages from resource consents through restrictions on the use of common resources (i.e. air, land and water). For example, 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. Furthermore, the time and costs to obtain resource consent for a new project may impose barriers to new industry entrants and confer an economic benefit on existing resource consent holders. However, the extent to which resource consent holders are able to benefit from the development of valuable internal capabilities such as environmental risk management systems, is not clear from the rather mixed findings of the environmental performance literature.

In this study we consider the shareholder wealth implications of capital asset expenditures that must comply with environmental regulations. The investment opportunities theory suggests that managers are able to maximise the market value of the firm by undertaking positive NPV projects (Miller & Modigliani, 1961). Much of the empirical research on capital expenditures seeks to test aspects of this theory.

One thread of literature has investigated the proposition that strategic capital expenditures may be a source of competitive advantage to firms (Porter, 1979). Theoretical research suggests that first movers who invest in cost-reducing technology earn economic rents during the early stages of the industry life cycle (Jovanovic & Macdonald, 1994), and that firms gain competitive advantage by undertaking capital investments that assist them to develop expertise that creates valuable real options (Bernardo & Chowdhry, 2002). Some empirical studies find support for shareholder wealth maximisation through positive market reactions to news of increases in capital budgets (McConnell & Muscarella, 1985), particularly for firms with greater investment opportunities (Chung, Wright, & Charoenwong, 1998; Vogt, 1997). Similarly, positive valuation implications have been found for capital expenditure project announcements. Woolridge and Snow (1990) find that projects with expected investment horizons greater than 3 years experience the largest abnormal returns, 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, Ho and Shih (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, Lonie and Power (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. The results of studies of environmental regulatory events imply that environmental regulations may impose on firms direct costs that reduce their ability to earn economic profits. However the technology-innovation and environmental-investment literature suggest that early adopters and firms pursuing voluntary capital initiatives may be relatively advantaged by environmental regulatory requirements. Our research is able to fill a gap in the literature by assessing the role of expected resource consent compliance costs in explaining the shareholder wealth impact of capital expenditure announcements.

Our first hypothesis involves testing the validity of a constructed indicator of expected resource consent compliance costs. As the time to obtain consent increases, total compliance costs can be expected to increase (Office of the Associate Minister for the Environment, 2004). Hence the time to gain resource consent approval can be viewed as an indicator of resource consent compliance costs. Accordingly, a composite variable *ETC*, the expected time for a project to gain consent, and a related dummy variable *ETCDUM*, are constructed to take account of investors' expectations of each project's consent compliance costs at the time of the project initiation announcement. The rationale for, and details of the construction of the *ETC* and *ETCDUM* estimates are presented later in section 4.2 of this paper, but if the *ETCDUM* variable is an appropriate indicator of the time for a given project to gain resource consent, then as postulated in hypothesis *H1*, it is expected that there will be a positive relationship between *ETCDUM* and the actual time to either obtain resource consent approval or abandon the project (*TCA*).

H1: The expected time for a project to gain resource consent 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 (as proxied by the *ETCDUM* variable) allow them to create sustainable advantages that are not easily replicated by their competitors, 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.

4. Data and methodology

4.1 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 i-Search and IRG Deep Archive databases between January 1991 and August 2007. The cut-off date was set to avoid the possible influence on results of a September 2007 government announcement of the planned introduction of an emissions trading scheme. Projects are only considered for inclusion in the sample if a resource consent to undertake or operate a project is either required or already possessed. Keyword search terms include ‘purchase’, ‘develop’, ‘development’, ‘acquire’ and ‘acquisition’ together with ‘consent’, ‘notify’, ‘non-notified’, and variations thereof. Capital expenditure projects are defined as the acquisition or construction of new plant and equipment, and the upgrade of existing tangible capital assets. 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, Perote, & Pindado, 2003).

From the keyword search, 128 capital expenditure projects are able to be 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. 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 coverage of New Zealand newspaper, newswire and magazine reports. As Newztext Plus has limited coverage prior to 1995, the Factiva media database is also used to identify pre-1995 announcements. Similar to Burton, Lonie & Power (1999), the earliest of the NZX i-search or media reporting date is designated as the event day, focusing upon a two-day (0,+1) event window. Stock exchange announcements after market close are deemed to arise on the next working day.

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 for which the initiation date can be clearly identified. Five projects are eliminated from consideration as no project or resource consent initiation dates can be identified. Second, no confounding events must occur within plus or minus two days of the announcement (-2,+2), resulting in 57 exclusions. Third, announcing firms' stock must have traded around the time of the announcement (-1,+1). Another 11 announcements are excluded as a result of this criterion. 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. Each of the 55 qualifying announcements was read to gather additional information such as the capital investment size, management's forecast of the expected time to gain consent and/or commence operations and whether or not the expenditure involves a joint venture. 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 (*ETC*) is constructed as explained in section 4.2 to estimate regulatory delay as the expected time (in months) to gain consent. Panel A of Table 1 gives details of the annual number of projects classified 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 the 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.

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 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 would still be implicit. 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. For the sample, the project news is first disseminated solely through the media for 26 projects (47%), solely through stock exchange releases for 16 projects (29%), and concurrently from both sources for 13 projects (24%). 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, a measure of stock market returns is needed. We construct an equal-weighted stock market index to avoid 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. 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 & Williams, 1977), leaving a broad index of 193 stocks. The announcement sample daily stock returns were then captured from the stock market index dataset.

To test hypothesis *H2a* concerning the effect of expected resource consent compliance costs on event-window abnormal returns, cross-sectional regression analysis is undertaken to test for differences between the two *ETCDUM* groups. Annual firm financial data including book value of total assets, book value of ordinary share equity and total liabilities is obtained from the NZX – Deep Archive Service. Market value of equity, market value of assets and industry classifications are sourced from Datastream. 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.

4.2 Measurement of expected time to consent

To test the shareholder wealth maximisation hypothesis in *H2a*, we need a measure of expected resource consent compliance costs at the time of project initiation. In a 2004 Cabinet briefing paper (Office of the Associate Minister for the Environment, 2004) proposing legislative changes that were later enacted in the Resource Management Amendment Act 2005, 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. Furthermore, the Ministry contends that it "is difficult to assess the influence that negative perceptions of the RMA have on investment certainty and decision making. However, 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.

Insert Table 2 about here

There is evidence to suggest that the time-varying costs associated with the time required to gain resource consent are particularly material. 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).

The foregoing analysis 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 based upon public information that investors could reasonably be expected to use at the project announcement date. Available public information contained in the media or stock exchange announcements may include management forecasts of the expected time to gain consent or commence operations, and past projects' actual times to gain consent. From this information we compile four forecast measures, summarised in Table 3, and detailed below. 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. The second measure is the historical firm-level median *TCA* (time to consent or abandon) based upon the past capital expenditure projects identified in the keyword search. For 87 out of the 128 capital expenditure projects, we are able to calculate the historical (actual) project-level *TCA* based upon the time elapsed from the reported initiation date to the subsequent resource consent granting/abandonment date. From this we prepare time-varying measures of the historical median *TCA* for each firm, which are updated each time a new consent is obtained. This results in historical firm-level *TCA* measures for 22 out of the 55 projects, with a mean (median) of 23.08 (12.30) months. Our 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 consent, it incorporates 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. For our last measure, the 87 projects for which the estimated historical project-level *TCA* is able to be calculated are used to compile a time-varying median *TCA* for each industry, updated as new consents are granted, or projects abandoned. This yields historical industry-level *TCA* measures for 41 of the 55 sample projects, with a mean (median) of 16.85 (13.81) months.

To evaluate which of the four forecast measures to incorporate into the *ETC* variable, we need to assess 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). Historical industry-level *TCA* is not associated with project-level *TCA* (ρ=0.0087, n=30) possibly because firm *TCA* varies considerably within each industry and industry sample sizes are very small. The correlations of the four component measures with the historical project-level *TCA* are repeated using logarithmic transformations, and result in no change in the rankings of the measures. Given the lack of correlation between historical project-level *TCA* and historical industry-level *TCA*, this measure is not used to construct the *ETC* variable.

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* (time to consent or abandon) for past capital expenditure projects. Seventeen estimates were gained from this measure.

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. The sample size for the correlation coefficient between this variable 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. The correlation is modestly positive at $\rho=0.1648$, however the *ETC* and historical project-level *TCA* variables are highly non-normal, so transformations are employed to improve consistency with the normality assumptions of the statistical tests. Taking the natural log of the 41 non-zero observations of the historical project-level *TCA* (*LNTCA*) improves the normality of this variable, as evidenced by the close proximity of the mean and median. To overcome the issue of non-normality for the *ETC* measure, a dummy variable is constructed by partitioning the sample at the median *ETC* of 11.2 months such that *ETCDUM* is equal to 0 if a given project *ETC* is below the sample median *ETC*, and equal to 1 if the project *ETC* is above the sample median *ETC*. The correlation coefficient between *ETCDUM* and *LNTCA* is $\rho=0.4148$, giving some credibility to proposition that the *ETCDUM* variable may be a reasonable indicator of project-level time to consent or abandon. Investor forecasts of project *TCA* are likely to contain substantial prediction error as does the *ETCDUM* measure calculated here, as actual project time to consent can vary widely within industries and even within firms. If the *ETCDUM* variable is an appropriate indicator of the expected time for a given project to gain resource consent, 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 we perform cross-sectional regression analysis using the following model.

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

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 "cost of approvals is not proportional to the business size" (Office of the Associate Minister for the Environment, 2004, p. 7). We control for size in two ways. First, following Chen and Ho (1997) we use the natural log of the market value of assets ($LNMVA_i$) as a proxy for firm size. Second, we consider the relative size of the project by calculating, where available, the dollar value of the investment divided by the book value of firm assets (INV/BVA_i) (Chen, 2006; Chen & Ho, 1997). 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 $LNTCA_i$ we use a dummy variable $REFORMDUM_i$ that takes the value of one for announcements in the post-reform period (after December 2002) and zero otherwise. We make no prediction regarding the sign of the $REFORMDUM_i$ coefficient.

4.3 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), we calculate abnormal returns as the difference between expected and observed market model returns over the event window. The market model is widely used in event studies and is appropriately specified with sample sizes as small as 50 (Brown & Warner, 1985; Corrado & Truong, 2008). We use the Scholes-Williams (1977) beta estimator to avoid the understatement of beta coefficients in the presence of infrequent trading.

Unreported analysis indicates that the sample distributions of security returns and abnormal returns violate the normality assumptions of parametric tests. Consequently we report results 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 & Zivney, 1992). For robustness, we also report the results using the Patell (1976) standardised abnormal return test, the Boehmer, Musumeci and Poulsen (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, and to consider other possible firm, industry or environmental variables that may influence our results, 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 ordinary least squares regression.

We estimate variants of the following cross-sectional regression model:

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

where CAR_i and $ETCDUM_i$ are described above, and the control variables are defined below. We initially include $ETCDUM_i$ as the sole independent variable in the regression equation, and then rerun the model to allow for additional control variables.

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 event period abnormal returns and firm size (Atiase, 1985). The results of studies of capital expenditure announcements tend to be consistent with this hypothesis (e.g. Chen, 2006; Chen & Ho, 1997; Jones et al., 2004). Accordingly, we follow Chen and Ho (1997) by measuring firm size as the natural log of the market value of firm assets ($LNMVA_i$) and expect to find a negative relationship with the stock market reaction.

If capital expenditure projects affect shareholder wealth, then relative project size may also be important. Consistent with Chen (2006) and Chen and Ho (1997) we measure project size as 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.

All the projects in the sample require (or already hold) resource consent in order to go ahead, however only two-thirds of the announcements make explicit mention of the consent. Although we expect that investors are sufficiently aware of New Zealand laws to understand that consent is required for major projects, it is possible that the additional information transmitted in announcements that discuss resource consents is valued by the market. Accordingly, we include in the model a dummy variable, $RCDUM_i$ equal to one when resource consent information is explicitly disclosed in the project announcement, and zero otherwise. We expect that if the resource consent disclosure is informative, then there will be a positive relationship 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. For the remaining 31 announcements, no prior project information is found from searching the stock exchange and media databases. 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 (Blacconiere & Patten, 1994; 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, we test for the impact of surprise project announcements using 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.

5. Empirical results

5.1 Measurement of expected time to consent

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 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 consent 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 consent. 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. Given that the time to gain resource consent approval can be viewed as an indicator of resource consent compliance costs (Office of the Associate Minister for the Environment, 2004), we use the constructed *ETCDUM* variable as an indicator of expected resource consent compliance costs.

Insert Table 4 about here

5.2 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 but consistent 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, and vary in significance from the 10% to the 1% levels. 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 U.S. (Woolridge & Snow, 1990), 0.86% in Singapore (Chen & Ho, 1997), 0.39% in the U.K. (Burton et al., 1999) and 0.30% in Korea (Kim, Lyn, Park, & Zychowicz, 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.

5.3 The influence of expected time to consent 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. 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 the announcement is anticipated by some investors. Only for the long *ETC* subsample is there any indication of a surprise element on the announcement day.

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 with a level of significance of 10%. When the explanatory variables for firm size (*LNMVA*) and relative project size (*INV/BVA*) are added in Models 3 and 4, the coefficient for *ETCDUM* remains significantly positive at the 10% level for *CAR* (0,+1), but becomes insignificant for *CAR* (0,+2). The negative coefficient (1% level) for *LNMVA* is consistent with similar studies suggesting that small firms experience greater information asymmetry (Chen, 2006; Chen & Ho, 1997). The sample size for Models 3 and 4 is relatively small as the project size was only disclosed for 32 out of the 46 announcements for which we have an *ETCDUM* estimate. 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 two further control variables, *RCDUM* and *DISCLOSDUM*. The t-statistics for *RCDUM* are positive and statistically significant at the 5% and 10% levels for the *CAR* (0,+1) and *CAR* (0,+2) models, respectively, indicating 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 coefficient on *DISCLOSDUM*, then the market reaction is negative. This is 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. The t-statistics for *ETCDUM* remain positive and significant at the 1% and 5% levels in the (0,+1) and (0,+2) windows.

In Models 7 and 8 of Table 8, we replace *ETCDUM* with the *ETC* variable and use rank regression to reduce the problem of non-normality. We transform the *ETC* and firm size variables to their ranked equivalents, *ETC RANK* and *MVA RANK*, and then apply conventional parametric generalised least squares regression. Both transformed variables are statistically significant at the 5% level or better.

In further (untabulated) analyses, we apply further robustness tests. We use dummy variables to test for possible effects from 2003 RMA reforms, the use of joint venture arrangements (Burton et al., 1999) and energy generation projects. We also check for the possible influences of growth opportunities and financial leverage on our regression results (Chen & Ho, 1997). The analysis is also repeated deleting a potentially important outlier. Overall, we find our 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. Our regression results indicate that the average values of capital expenditure projects with long expected times to consent exceed those with short expected times to consent by about 1.51% to 1.97% (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 1991 dollars is $13.0 to $17.0 million (relative to the average market value of equity of $860.8 million). In 2007 dollars, this translates to a net benefit of $18.0 to $23.4 million (relative to the average market value of equity of $1,189.3 million). These results suggest that on average, the marginal expected benefits of undertaking environmentally-sensitive projects with relatively long expected times to consent are substantially greater than the marginal expected compliance costs resulting from lengthy consent processing times.

The above findings also suggest that the stock market reaction to short-time-to consent project announcements is statistically no different from zero. For these projects, potential competitive advantages may be diminished if short consent 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.

6. Conclusion

This study investigates the role of environmental regulatory delay in explaining the capital market impact of New Zealand capital expenditures. Stock market reactions to project announcements are predicted to be positive if the shareholder wealth maximisation hypothesis holds. The corollary prediction is that projects with higher resource consent compliance costs create a sustainable advantage that makes them more valuable.

An indicator of expected resource consent compliance costs at the time of project initiation is constructed and found to be positively related to the actual time that it takes firms to either obtain resource consent approval or abandon the project. Using the expected compliance cost indicator to separate a sample of project announcements from 1992 to 2007 into two groups indicates that the market response is significantly positive when expected consent compliance costs are high and insignificant when they are low. Cross-sectional regression results confirm a significantly positive relationship between the expected time to consent and the market reaction to project announcements.

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 consent is long, which is consistent with the suggestion that firms that incur greater resource consent compliance costs create sustainable advantages that competitors cannot easily replicate. For an average firm undertaking a long time-to-consent project, the net benefit is estimated to be in the range of $18.0 to $23.4 million dollars (at 2007 prices). For these projects we suggest that the time delay for resource consent approval and high level of compliance costs incurred may allow firms to develop specialised capabilities and/or to deter industry competitors and new entrants, thereby increasing expected project NPVs. In contrast, announcements of new projects for which expected consent processing time is short generate a neutral stock market reaction. For these projects, we suggest that shorter consenting times and lower compliance costs are insufficient to allow firms to develop specialised capabilities or to impose barriers to entry, thereby driving expected projects' NPVs to zero.

Our results suggest that in a small, open economy, environmental compliance costs may have the effect of enhancing the competitive position of those firms that undertake major capital expenditures. These findings also imply that if legislators are able to reduce environmental regulatory delays associated with capital expenditures through further enhancements to the RMA legislation, then the opportunity for firms to earn economic profits may be diminished.

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.

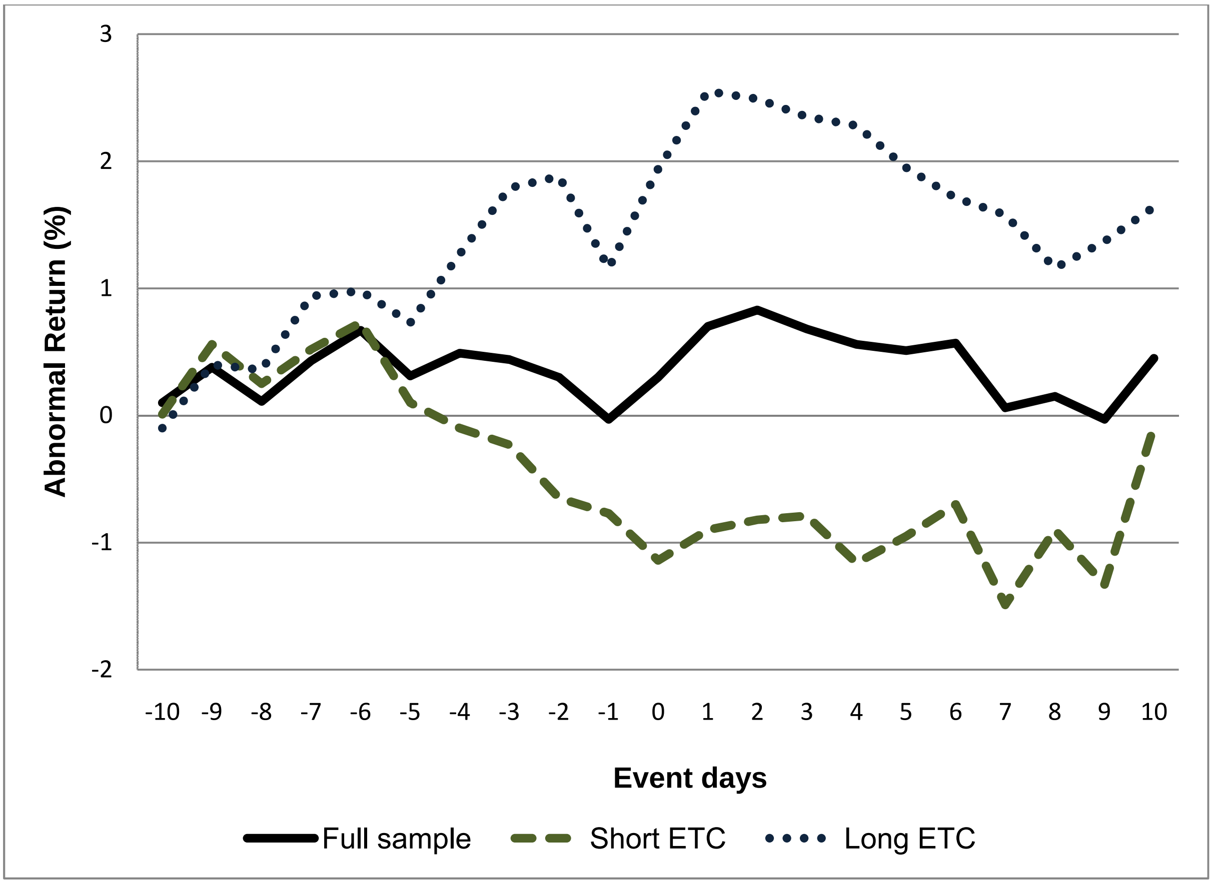

Table 1 Project sample analysis

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 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, 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, 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 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), management forecast months for the project to become operational and the historical industry-level TCA. 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 the first three component variables that are most highly 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, b, and c denote statistical significance at the 1%, 5% and 10% levels.

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 industry-level TCA	41	16.85	13.81	9.59	6.76	47.05	0.0087 (30)
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

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 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, b, and c denote statistical significance at the 1%, 5% and 10% levels.

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)[b]
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 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 ,b and c denote statistical significance at the 1%, 5% and 10% levels 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[c]	-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[c]	2.08[b]	1.99[b]	1.89[c]
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[c]	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[c]	-1.59	-1.58	-1.82[c]
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[c]	2.26[b]	2.18[b]	1.92[c]
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 ,b and c denote statistical significance at the 1%, 5% and 10% levels 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[c]	-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[c]	-1.90[c]	-1.86[c]	-2.11[b]
0	0.0077	0.0005	0.52	0.3948	0.2968	1.33	1.95[c]	1.65[c]	1.33
1	0.0062	0.0030	0.61	0.4956	0.2968	1.67	2.33[b]	1.75[c]	1.74[c]
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, b and c denote statistical significance at the 1%, 5% and 10% levels.

	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[c]	-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[c]	-1.42

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

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 when resource consent information is explicitly disclosed in the project announcements and 0 otherwise, DISCLOSDUM equals 1 for first disclosures, and 0 otherwise and ENERGDUM equals 1 if the announcement relates to an energy generation projects, and 0 otherwise. In Models 7 and 8, the analyses are performed using rank regression. ETCDUM is first replaced with the ETC variable which is then transformed to its ranked equivalent ETCRANK. Similarly, the rank of market value of equity, MVARANK, is used in place of LNMVA. 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 ,b and c denote statistical significance at the 1%, 5% and 10% levels.

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)	CAR(0,+1) Model 7 (n=46)	CAR(0,+2) Model 8 (n=46)
Constant	n/a	-0.0012 (-0.29)	-0.0003 (-0.06)	0.2455 (3.79)[a]	0.1810 (2.46)[b]	0.2749 (4.82)[a]	0.2183 (4.73)[a]	0.0074 (0.78)	0.0063 (0.56)
ETCDUM	+	0.0151 (1.89)[c]	0.0136 (1.92)[c]	0.0149 (1.97)[c]	0.0120 (1.62)	0.0197 (2.55)[a]	0.0159 (2.42)[b]		
ETC RANK								0.0002 (2.13)[b]	0.0004 (2.38)[b]
LNMVA	-			-0.0121 (-3.83)[a]	-0.0087 (-2.49)[b]	-0.014 (-4.79)[a]	-0.0109 (-4.85)[a]		
MVA RANK								-0.0000 (-5.56)[a]	-0.0000 (-4.35)[a]
INV/BVA	+			-0.0002 (-0.01)	0.0008 (0.02)				
RCDUM	+					0.0173 (2.01)[b]	0.0122 (1.69)[c]	0.0112 (1.47)	0.0090 (1.16)
DISCLOSDUM	n/a					-0.0125 (-1.89)[c]	-0.0151 (-2.37)[b]	-0.0112 (-1.54)	-0.0123 (-1.71)[c]
Adjusted R^2		0.051	0.053	0.247	0.173	0.358	0.315	0.275	0.255

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