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Abstract

We investigate the applicability of the present-value asset pricing model to fishing quota markets by applying instrumental variable panel data estimation techniques to 15 years of market transactions from New Zealand's individual fishing quota market. In addition to the influence of current fishing rents (as measured by lease prices), we explore the effect of market interest rates, risk, and expected changes in future rents on quota asset prices. Controlling for these other factors, the results support a fairly simple relationship between quota asset and contemporaneous lease prices. Consistent with theoretical expectations, the results indicate that quota asset prices are positively related to declines in interest rates, lower levels of risk, expected increases in future fish prices, and expected cost reductions from rationalization under the quota system. However, the magnitude of some interrelationships is muted relative to what theory suggests, possibly due to measurement error.

Key Words: tradable permits, individual transferable fishing quota, asset pricing, fisheries, policy

JEL Classification Numbers: Q22, Q28, D40, L10

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Introduction

In December 2004, the U.S. government released an Ocean Action Plan that encourages the fishery management councils—regional governing bodies that set total catches and other regulations—to adopt market-based systems for fisheries management. Individual fishing quotas, in which the total catch is capped and shares of the catch are allocated, are one such system. An individual transferable quota (ITQ) system results when transfer of the shares is permitted, and the least efficient vessels will find it more profitable to sell their quota rather than fish it. Over time, this should both reduce excess capacity and increase the efficiency of vessels operating in the fishery.

For ITQs to deliver an efficient solution to the common pool problem in practice, it is critical that quota markets are competitive and convey appropriate price signals. Price signals sent through the quota market are therefore an essential source of information on the expected profitability of fishing and an important criterion for decisions to enter, exit, expand, or contract individual fishing activity. Quota prices also send signals to policymakers about the economic and biological health of a fishery. For example, Arnason (1990) showed that under the assumption of competitive markets, monitoring the effect of changing the total allowable catch (TAC) on quota prices could be used to determine the optimal TAC.

In a previous study, Newell *et al.* (2005) investigate the performance of ITQ markets using the most comprehensive dataset gathered to date for the largest system of its kind in the world. The panel dataset from New Zealand covers 15 years of transactions across the 33 species that were in the program as of 1998 and includes price and quantity data on transactions in more than 150 fishing quota markets. Markets exist in New Zealand both for selling the

perpetual right to a share of a stock's TAC, as well as for leases of that right to catch a given tonnage in a particular year.[1] Newell *et al.* (2005) found that market activity appears sufficiently high to support a reasonably competitive market for most of the major quota species and that price dispersion has decreased over time. Investigating the asset and lease markets *separately*, they find evidence of economically rational behavior in each of the quota markets and their results show an increase in quota asset prices, consistent with increased profitability.

We extend the analysis of Newell *et al.* (2005) by econometrically examining the relationship *between* the annual lease and sale of the perpetual quota asset markets.[2] With competitive markets, rational asset pricing theory suggests that the price of an income-producing asset in period t, p_t, should be determined by the real per-period profits from the asset, π_t, and the real discount rate (r):

(1)
$$ p_t = \sum_{j=0}^{\infty} \frac{\mathrm{E}_t(\pi_{t+j})}{\prod_{k=0}^{j}(1+\mathrm{E}_t(r_{t+k}))}, $$

where $\mathrm{E}(\cdot)$ is the expectations operator. In our setting, equation (1) states that the current quota asset price should be equal to the present discounted value of all future expected earnings, where the lease prices represent the annual flow of profits from holding quotas. The price of the quota asset, therefore, will vary across fish stocks and over time based on changes in expected future lease prices or changes in the expected discount rate over time.

Under the simplifying assumption that expected lease prices and discount rates remain constant in the future, the price of the asset would simply equal the lease price divided by the discount rate, or $p_t = \pi_t / r_t$. The expected rate of return from holding fishing quota (or dividend-price ratio) would be equal to π_t / p_t. Figure 1 supports the basic structure of such a relationship in New Zealand fishing quotas, with the dividend-price ratio tracking both the level and the trend in New Zealand short-term interest rates over the sample period. For example, at the same time

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[1] Virtually all leases are for a period of one year or less.

[2] Batstone and Sharpe (1999) investigate the relationship between fishing quota and lease prices and changes in the total allowable catch for the New Zealand red snapper fishery (region 1) and find support for the relationship proposed by Arnason (1990). Other related research in fisheries includes Karpoff (1984a) and Huppert *et al.* (1996), which look at the relationship between license prices and fishery rents in Alaska salmon fisheries.

the dividend-price ratio fell by about half from 13 percent to 7 percent, the interest rate as measured by New Zealand Treasury bills fell from 10 percent to about 4 percent in real terms. Overall, the quota dividend-price ratio is about 2–3 percent higher than the risk-free rate on average. Figure 2 likewise suggests a close, relatively linear association between asset and lease prices (in logs). The level of the average asset price is also approximately 10 times the lease price over the sample period, roughly equal to the present value of a perpetuity discounted at 10 percent.

Figure 1 also shows that there is considerable cross-sectional variation in the dividend-price ratio across fish stocks markets, where the upper and lower plus signs represent the 25th and 75th percentiles around the median. Why might such variation exist? One reason could be that if fishers are risk averse, they might prefer fish stocks with lower variance, other things equal. This effect is consistent with a higher discount rate, or higher required rate of return for riskier stocks. Such volatility could be associated with natural variation in stock abundance and economic variability in costs and fish prices. Another explanation could be differences in the expected growth rate of profits over time (Melichar 1979), possibly due to differences in output price growth, changes in fish populations, or other factors affecting costs such as cost rationalization due to quota trading.

Using panel data econometric techniques on an updated Newell *et al*. (2005) dataset, we estimate models that relate the asset price of quotas to their annual lease (or rental) price and observed determinants of the growth rate and volatility of rents. Within this framework, we explore the relationship between asset and lease prices, as well as whether differences in asset prices are due to differential risks associated with holding quotas across fish stocks and/or different expected growth rates in fishery rents in those stocks. These data are uniquely qualified to address these questions, because of the relatively long time series, breadth of markets, and cross-sectional heterogeneity, as the market characteristics are diverse across both economic and ecological dimensions (see Table 1 for a list of species included). For example, in 2000 the export value of these species range from about NZ$700 per ton for jack mackerel to about NZ$40,000 per ton for rock lobster.[3]

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[3] Throughout this paper, monetary values are year 2000 New Zealand dollars, which are typically worth about half a U.S. dollar. Tons are metric tons.

Consistent with asset pricing theory, we find a statistically (and economically) significant relationship between asset prices and contemporaneous lease prices. Stocks with a higher degree of biological volatility tend to have lower asset prices, and stocks that have rising returns or falling costs from fishing are found to have higher asset prices, *ceteris paribus*. Taken together, these results suggest that the price signals generated by the ITQ system are a good indication of the future profitability of individual fishing quota stocks. The magnitude of some interrelationships is muted relative to what the theory suggests, possibly due to measurement error.

In the next section, we describe the design of the ITQ system in New Zealand, paying particular attention to market characteristics. This is followed by a selected review of the literature modeling asset prices and dividends. We then discuss the empirical specification, data sources, time-series properties of the data, estimation approach, and results, before we conclude by summarizing our findings.

New Zealand's Individual Transferable Quota (ITQ) System

After several years of consultation with industry, the government of New Zealand passed the Fisheries Amendment Act of 1986, creating a national ITQ system.[4] The system initially covered 17 inshore species and 9 offshore species, which together expanded to a total of 45 species by 2000. Under the system, the New Zealand Exclusive Economic Zone (EEZ) is geographically delineated into quota management regions for each species based on the location of major fish populations. Rights for catching fish are defined in terms of fish stocks that correspond to a specific species taken from a particular quota management region. In 2000, the total number of fishing quota markets stood at 275, ranging from 1 for the species hoki to 11 for abalone. As of the mid-1990s, the species managed under the ITQ system accounted for more than 85 percent of the total commercial catch taken from New Zealand's EEZ and from our calculations had an estimated market capitalization of about NZ$3 billion.

The New Zealand Ministry of Fisheries sets an annual total allowable catch for each fish stock based on an intertemporal biological assessment (including the prior year's catch level) and other relevant environmental, social, and economic factors. The TACs are typically set with a

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[4] For further history and institutional detail, see Batstone and Sharp (1999), Yandle (2001), Newell *et al.* (2005), and the references cited therein.

goal of moving the fish population toward a level that will support the largest possible annual catch (*i.e.*, maximum sustainable yield), after an allowance for recreational and other noncommercial fishing. Individual quotas were initially allocated to fishermen free of charge as fixed annual tonnages in perpetuity based on their average catch level over two of the years spanning 1982-1984. Beginning with the 1990 fishing year, however, the government switched from quota rights based on fixed tonnages to quotas denominated as a share of the TAC. Compliance and enforcement is undertaken through a detailed set of reporting procedures that track the flow of fish from a vessel to a licensed fish receiver (on land) to export records, along with an at-sea surveillance program including on-board observers.

   Given the uncertainty around the quantity and composition of catch, a fisherman's quota holdings represent a mix of ex ante and ex post leases, as well as asset purchases and sales to cover actual catch.[5] Whether ex ante or ex post transactions, fishing quotas are generally tradable only within the same fish stock, and not across regions or species or years, although there have been some minor exceptions.[6] The quota rights can be broken up and sold in smaller quantities and any amount may be leased and subleased any number of times. There are also legislative limits on aggregation for particular stocks and regions, and limitations on foreign quota holdings.[7]

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[5] Although there are no official statistics, the general belief among government officials and quota brokers is that a majority of the transactions between small and medium-sized quota owners are handled through brokers. Larger companies, on the other hand, typically have quota managers on staff and engage in bilateral trades with other large companies. A brokerage fee between 1 percent and 3 percent of the total value of the trade to be paid by the seller is standard.

[6] During the time period of our analysis, in addition to the lease and asset markets, fishers had a number of ways within a 30 day window after they landed their catch to balance their quota holdings and catches. First, a "by-catch trade-off exemption" allowed fishers who incidentally take non-target fish to offset the catch by using quota from a predetermined list of target species. Second, quota owners could carry forward to or borrow from the next year up to 10 percent of their quota (although not leases). A third option was to enter into a non-monetary agreement to fish against another's quota. Alternatively, a fisher could surrender the catch to the government or pay a "deemed value," which is set based on the nominal port price to discourage discarding of catch at sea and targeting stocks without sufficient quota (Annala 1996). These rules have changed somewhat since October 1, 2001, when annual quota leases were supplanted by sales of "Annual Catch Entitlements" or ACEs, which are issued annually by the government equal to each quota owner's annual quota allocation.

[7] Initially, the aggregation limits were on holding quota. Substantial changes were written into the 1996 Fisheries Act, including changing the limits on holdings to ownership levels, and limits for particular species and region combinations.

From 1986 to 2000, Newell *et al*. (2005) find that the quota markets are active,[8] with about 140,000 leases and 23,000 quota asset sales occurring between economically distinct private entities—an annual average of about 9,300 leases and 1,500 asset sales.[9] The annual number of leases has risen 10-fold during this period, and the median percentage of quotas leased in these markets has risen consistently, from 9 percent in 1987 to 44 percent in 2000. At the same time, the total number of quota asset sales declined from a high of about 3,200 sales in 1986 (when initial quota allocations for most species took place), leveling off to around 1,000 sales in the late 1990s. The median shows a similar decline, with the percentage sold being as high as 23 percent at the start of the program, gradually decreasing in subsequent years to around 5 percent of total outstanding quotas per year in the late 1990s. This pattern of asset sales is consistent with a period of rationalization and reallocation proximate to the initial allocation of quotas, with sales activity decreasing after the less profitable producers have exited.

Modeling Asset Prices and Dividends

The literature investigating asset prices and their relationship to dividends and other factors (*e.g.*, price-earnings ratio and firm size) is extensive. A thorough literature review is therefore beyond the scope of this paper, and interested readers should consult Campbell *et al*. (1997) or the review articles by LeRoy (1989), Fama (1991), and Campbell (2000). Our discussion of relevant literature is focused on studies investigating agricultural land prices and farming rents (*e.g.*, Melichar 1979; Alston 1986; Falk 1991;Clark *et al*. 1993; Just and Mirinowski 1993), agricultural production quotas (*e.g.*, Barichello 1996; Wilson and Sumner 2004), fishing rents and license prices (Karpoff 1984a, 1984b, and 1985; Huppert *et al*. 1996),[10]

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[8] Although the typical ITQ market exhibits a reasonably high degree of activity, some individual quota markets are thin. Quota markets with low activity tend to be of low economic importance in the size and value of the catch.

[9] They also find that about 22 percent of quota owners took part in a market transaction in the first full year of the program, increasing steadily to around 70 percent by 2000.

[10] Karpoff (1984b) tested and found evidence for demand and supply effects on license prices due to a shift in a government loan program during the course of the limited-entry system. In a related paper, Karpoff (1985) found that license prices are mainly determined by pecuniary factors (rents) but after identifying the continued presence of low-income fishermen, he argues that their presence is weak evidence for the existence of non-pecuniary benefits in license prices.

and fishing quota prices (Batstone and Sharp 2003).[11] The methodological approaches applied are closely related across all these asset price analyses.

Simplifying equation (1) under the assumption that the expected discount rate follows a martingale process[12] yields

(2)
$$p_t = \sum_{s=0}^{\infty} (\frac{1}{1+r_t})^{s+1} E_t(\pi_{t+s}).$$

This illustrates how the asset price is dependent on the expected future stream of earnings, so that information available at time t along with type of expectation process is important in modeling the relationship between asset prices and dividends. For example, if one assumes that expected future earnings are constant, then $E_t(\pi_{t+s}) = E_t(\pi)$. Huppert *et al.* (1995) model an adaptive expectations process where $E_t(\pi) = \beta \pi_{t-1} + (1-\beta) E_{t-1}(\pi)$ with $\beta = [0,1]$, and Karpoff (1984b) models a myopic process where $\beta = 1$. Wilson and Sumner (2004) model a second-order adaptive expectation model when investigating California dairy quota prices. Just and Miranowski (1993) test myopic, adaptive, and rational expectation regimes and find that farmland price data support myopic expectations.[13] Falk (1991) finds a similar result.

If future profits (lease prices) grow at a constant rate g, then $\pi_t = (1+g)\pi_{t-1} + \varepsilon_t$, where ε_t is a white noise error term. Taking expectations and solving equation (2) forward in time with $g < r$, the asset price follows[14]

(3)
$$p_t = \frac{\pi_t}{r_t - g}.$$

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[11] There is a significant literature investigating aspects of market efficiency in a variety of markets, including those for financial assets (Cochrane 2001; Fama 1998), housing (Case and Shiller 1989), art (Pesando 1993), orange juice concentrate (Roll 1984), and natural resources (*e.g.*, lease and sale of land, oil fields, and forest tracts). Recently, McGough *et al.* (2004) investigate the implications of a rational expectations equilibrium in timber markets on the time series properties of timber prices.

[12] If one assumes that discount rates follow a martingale process, then the best predictor of expected future discount rate at time t is the current rate, *i.e.* $E_t(r_{t+1})=r_t$ (LeRoy 1989). This is supported empirically by econometric analyses of interest rates. For a more general analysis of time-varying rates, see Chapter 7 of Campbell *et al.* (1997).

[13] Orazem and Miranowski (1986) provide an empirical strategy for testing competing hypotheses of expectations regimes when direct measures of expectations are unavailable. Applied to farm acreage allocation decisions as a function of expected commodity prices, it yielded little evidence for favoring any of the three regimes they tested.

[14] This equation differs from that of Melichar (1979) because here π is expressed in terms of the end-of-period value of returns from the asset, not the pre-period value.

Equation (3) is the dynamic "Gordon growth model" (Campbell *et al.* 1997) that forms the basis of the majority of studies on the relationship between asset prices and dividends. Batstone and Sharp (2003) use equation (3) with $g = 0$ as the basis for their analysis of the effects of TAC changes on asset and lease prices (see footnote 2).

Due to a divergence between simple, present-value relationships and empirical observations of agricultural land prices and rents during the 1970s and 1980s, a number of authors have extended the basic structure to include other factors, such as taxes (*e.g.*, Robison *et al.* 1985; Alston 1986), changes in risks (Barry 1980), and credit market constraints (Shalit and Schmitz 1982). Instead of investigating the many factors separately, Just and Miranowski (1993) develop a detailed structural model of the determinants of asset prices, which is a function of inflation, taxes, credit market imperfections, transaction costs, and risk aversion.[15]

Others have focused on estimating a reduced form that is consistent with equation (2). For example, Burt (1986) argues that movements in asset prices may occur because of continued adjustment to past changes in returns, implying that the price does not necessarily adjust instantaneously to changes in expected future returns. In addition, expectations of future rents may be based on past, as well as current, values of π_t. He approximated the effect of both sources of dynamic behavior by assuming a multiplicative distributed lag specification for π_t with a restriction that the lag coefficients sum to unity so that the dynamic representation would converge to the long-run equilibrium of equation (3) with no expected growth in rents.

In the next section, we generalize equation (3) to account for the unique features of an individual fishing quota market.

Empirical Analysis of Fishing Quota Asset Prices

Empirical Model

Our empirical assessment of the relationship between quota asset prices and expected future profits from fishing quota is based directly on the dynamic Gordon growth model (equation (3)). Within this framework, we explore possible explanations for the heterogeneity in quota asset prices across the different fish quota markets, as illustrated in Figure 1. Potential

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[15] In the absence of these complications, their price equation reduces to an expression that is identical to equation (3) with no growth.

reasons for the heterogeneity include different growth rates of profits due to expected changes in revenues or costs or because fish stocks are associated with different risk premia.

It is straightforward to allow for different asset prices, profits, and expected growth rates of profits across fish stocks by indexing each by an *ij* combination indicating a different fishing quota market, where species are denoted by *i* and regions by *j*. To investigate different risk premia, we follow the methods employed in Alston (1986) and Cochrane (1992) by decomposing the discount rate into a real market interest rate ($\tilde{r}_t$) and an asset-specific risk premium (θ_{ij}). Formally, we have

(4)
$$p_{ij,t} = \frac{\pi_{ij,t}}{(\tilde{r}_t + \theta_{ij} - g_{ij})} \ .$$

In fishing quota markets, a major difference in risks stems from ecological volatility, where some fish stocks have more variable populations from one year to the next. The greater fluctuations in population abundance could lead to greater cost and harvest uncertainty, as searching costs depend on the stock size and location.

In our setting, another important issue arises when considering the application of equation (4) which, for simplicity, assumes continuous growth into the indefinite future. In particular, fishing quota markets are created to address the "tragedy of the commons", and our analysis includes a period over which there was a market-based transition away from regulated open access conditions. Typically, when quota markets are created, fishing capital and labor inputs are distorted and fish populations are depleted due to years of operating under regulated open access conditions. An implication of this is that there will likely be a divergence between the current lease-asset ratio and the longer-term equilibrium, at least early on in the market, because at that time the contemporaneous lease price is not a good indicator of future profitability. This means that the asset price of a stock anticipating rationalization would initially be relatively high compared to its lease price. This divergence would decline over time as the stock achieved its anticipated profit increases and higher lease prices. Figure 1 suggests support for this hypothesis, as the difference between the 25th and 75th percentiles follows a downward trend.

Why might the divergence decrease over time? Initially, trades of the perpetual right to fish will occur as high-cost fishers find it profitable to sell their quota rather than fish it. The

gains from trade and elimination of excess fishing capital will result in expected cost savings. In addition, if the TACs are set to allow stock recovery, then the gains due to stock rebuilding will also be incorporated in the expectation of future costs.[16] The ability to time fishing trips to higher product prices rather than being forced to operate in short seasons, along with the shift from maximizing quantity to quality will also feature in near term expectations of future revenue growth.[17] These effects will likely dissipate over time as the potential gains are realized, where the rate of dissipation is an empirical question.

We modify equation (4) to account for these transitory effects by including a multiplicatively separable function (Ψ) representing the transition associated with ITQs:

$$(5) \qquad p_{ij,t} = \frac{\pi_{ij,t}}{\tilde{r}_t + \theta_{ij} - g_{ij}} \Psi(\cdot).$$

We expect $\Psi(\cdot)$ to be greater than one, because asset prices in ITQ markets will be initially above the long-run relationships due to short-run expected profitability gains. Furthermore, we expect it to be larger for stocks with greater short-term gains, but to be decreasing over time, as asset prices should converge to the long-run relationship after some interval of time, holding everything else equal. The arguments of $\Psi(\cdot)$ can include, for example, time since the market was created, and variables that represent gains from trade and fish stock recovery.

Empirical Specification and Data

After adding and subtracting 1 in the denominator of equation (5) (see footnote 21), we take a logarithmic approximation. We also approximate $\ln \Psi(\cdot)$ by $\beta_5 s_{ijy} + \beta_6 a_{ij} + \beta_7 a_{ij} t_y$, where s is a measure of expected future cost declines due to reallocation of fishing effort through trading, a indicates the extent of expected future cost declines to increases in fish stock abundance, and t is an annual time index.[18] Specifically, the relationship we bring to the quota asset price data is

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[16] In many fisheries, the cost function is likely to be stock-dependent, so that costs increase as the fish stock size falls, as it becomes harder to find the fish (*i.e.* searching costs increase).

[17] For example, since the introduction in 1994 of an ITQ system in the Alaskan halibut fishery, the season length has grown from two 24-hour openings to more than 200 days. The flexibility to time fishing trips when port prices are higher, and the elimination of large supply gluts of fresh product, have resulted in increases in price per pound of more than 40 percent (Casey *et al.* 1995). The focus on quality is also evident in New Zealand, where fishermen have changed catching methods in the red snapper fishery in order to sell their catch on the highly profitable Japanese live fish market (Dewees 1998).

[18] Ideally $\lim_{t \to \infty} \ln \Psi(\cdot) = 1$, but this would require a functional form necessitating nonlinear estimation in an instrumental variables panel data context. We have therefore opted for a linear approximation.

$$\begin{aligned}(6) \quad \ln p_{ijmy} &= \beta_1 \ln \pi_{ijmy} + \beta_2 \ln(1+\tilde{r}_{my}) + \beta_3 \ln \theta_i + \beta_4 \ln(1+g_i) + \beta_5 s_{ijy} + \beta_6 a_{ij} + \beta_7 a_{ij} t_y \\ &\quad + \beta_8 d_i + \alpha_m + \alpha_y + \nu_{ij} + \varepsilon_{ijmy} \end{aligned},$$

where p is the quarterly average quota asset price, π is the contemporaneous quota lease price (as a measure of the annual profits from fishing), $\tilde{r}$ is the real interest rate, $\ln\theta$ is proxied by each species natural mortality rate (a measure of risk), and g is proxied by a measure of expected future growth in the output price of fish species i. We also include a dummy variable (d) for shellfish stocks (*i.e.*, abalone, rock lobster, and scallops), a set of quarterly fixed effects (α_m), a set of yearly fixed effects (α_y), a fish-stock-specific effect (ν) whose specification varies depending on the estimation approach (*e.g.*, fixed or random effects), and an independently identically distributed error term ε. Species are denoted by the subscript i and regions by j, so that each ij combination indexes a different fishing quota market. Time is indexed by quarter m of year y.

The model and accompanying discussion above imply the following hypotheses for the model: $\beta_1 > 0$, $\beta_2 < 0$, $\beta_3 < 0$, $\beta_4 > 0$, $\beta_5 > 0$, $\beta_6 > 0$, and $\beta_7 < 0$. Strict interpretation of the logarithmic approximation given by equation (6) further implies the following hypotheses about the specific magnitudes of certain coefficients: $\beta_1 = 1$, $\beta_2 \approx -(1+r)/(r+\theta-g)$, $\beta_3 \approx -\theta/(r+\theta-g)$, and $\beta_4 \approx (1+g)/(r+\theta-g)$, where each of the variables in these formulae is taken to be its mean value (the point of approximation). We do not impose these as restrictions, but rather consider them when interpreting the findings below.

We estimate equation (6) using the comprehensive panel dataset described in detail in Newell *et al.* (2005), which was constructed using information from New Zealand government agencies and other sources. We include 152 fish stocks representing 32 different species that had entered the New Zealand ITQ system by 1998. The data cover 15 years from the 1987-1988 fishing year until the end of the 2000-2001 fishing year. All monetary figures were adjusted for inflation to year 2000 New Zealand dollars, using the producer price index (PPI) from Statistics New Zealand. Table 2 gives descriptive statistics for the 4,120 observations comprising the estimation sample; the included variables exhibit a large degree of variation.

As described above, the quota asset and lease prices are quarterly averages for each species-region specific fish stock quota market, based on more than 140,000 underlying lease transactions and more than 23,000 asset transactions.[19] The real market interest rate, $\tilde{r}$, is the 90-day New Zealand Treasury bill rate, adjusted for inflation using the New Zealand CPI. As a measure of variation in the risk premium across species, lnθ, we use each species' natural mortality rate. Species with higher mortality rates have population sizes that are typically more variable due to fewer age classes, which we argue leads to increasingly greater uncertainty in the amount of fish likely to be caught with a given level of effort. As a consequence, there is greater uncertainty in the profits from fishing high-mortality species, and we would therefore expect higher mortality rates to have a negative effect on quota asset prices.[20] We base g on the historic growth rate in output prices, where output prices are based on the export price per greenweight ton using data from Statistics New Zealand over the period 1986–2001, deflated using the NZ PPI (see Newell *et al.* 2005).[21]

Empirically, the components of the approximation to the $\Psi(\cdot)$ function are as follows. To represent expected future profit increases due to reallocation of fishing effort through trading, s, we use the annual percentage of quota assets sold for each fish stock, normalized by dividing by each stock's average percentage sold. The hypothesis is that reallocation of quota assets is an indication of expected future profits from that trade, most likely through cost reductions. Improvements in profits through cost reductions can also occur as a result of improvements in fish stock abundance and associated increases in the catch-per-unit-effort. We represent this feature using a dummy variable, a, that indicates whether each stock faced significant reductions

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[19] Prices were available for 151,835 leases and 25,210 sales. About 30 percent of lease and 25 percent of sale observations that did not represent reliable market transactions were omitted. For more information on how these prices were identified, see Newell *et al.* (2005).

[20] The New Zealand Ministry of Fisheries uses the mortality rate to construct a measure of natural variability that is factored into the setting of the TAC (Annala *et al.* 2000). The assumption is that a stock with higher natural mortality will have fewer age classes and therefore have greater fluctuations in biomass.

[21] We estimate the output price growth rate independently for each species based on a first-order autoregressive model of the log fish price, including a time trend, quarterly (seasonal) effects, and a constant term. The estimates for a small number of species are negative and to avoid taking a logarithm of a negative number, we add and subtract 1 in the denominator of equation (5). Another option would be to directly estimate the growth rate in lease prices, but this introduces econometric issues due to the endogeneity of lease prices.

upon implementation of the ITQ program.[22] We expect that fisheries plagued by excess capacity and overfishing prior to the implementation of the ITQ system and also faced significant reductions in allowable catch at the outset of the ITQ program would experience greater increases in profitability through stock rebuilding and cost rationalization than fish stocks without a high degree of overfishing, everything else being equal. Thus, we would expect the coefficient on a to be positive, indicating that for a given lease price, the asset price will be higher for stocks with fish stock rebuilding plans in place.

Over time, however, the gains from such improvements should be realized, implying that future gains will be lower. We capture this effect by interacting a with a time trend, hypothesizing that over time the lease price will rise as stocks improve, and the effect on the asset price of additional future gains will diminish. Under these conditions, we would expect the coefficient on the interaction of a and t to be negative

Estimation Approach

Time-Series Properties of Data

Before considering estimation of equation (6), it is essential to determine the time series properties of the asset and lease price series. If either one or both of the series are non-stationary, then standard regression techniques will be susceptible to the problem of spurious regression. If both are non-stationary, however, a cointegrating relationship may exist between the two series (see, for example, Campbell and Shiller 1987).

While testing for unit roots in panels is a relatively new enterprise, there are several tests available to researchers (see Banjeree 1999 for more information on the tests). In this paper, we employ three tests, all of which can be thought of as panel data extensions or pooled versions of the Dickey-Fuller test (or Augmented Dickey-Fuller when lags are included). The null hypothesis is that the series are non-stationary of order one. In individual series, it is well-known that the power of unit root tests is low. Panel unit-root tests, however, have been shown in Monte Carlo analysis to have significantly higher power (Banjeree 1999; Levin *et al.* 2002).

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[22] We classified fish stocks as to whether they faced significant initial catch reductions under ITQs by using historical information on catch rates, TAC levels, and references in the literature (see supplementary material in Newell *et al.* (2005) for more information). The following 33 fish stocks were so classified: CRA1-5, CRA7-8, BNS2, ELE3-5, JDO1, MOK1-3, ORH2B, SCH1-3, SCH5, SCH7-8 SKI3, SNA1-2, SNA8, SPO1-3, SPO7-8, TRE1, HPB2-3.

First, we use the test by Levin, Lin, and Chu (LLC, 2002), which was first published as a working paper in 1993 and was formerly known as Levin and Lin's test. This test assumes that each fish stock shares the same AR(1) coefficient, but it does allow for individual fish stock effects, time effects and time trends, and serial correlation in the errors. An advantage of this test over the test proposed by Hadri (2000), for example, is that it is better suited for medium-sized panels similar to ours. We also use the test developed by Im, Pesaran, and Shin (IPS, 2003). This test is similar in spirit to that of Levin and Lin (1993), but the alternative hypothesis allows for heterogeneity in the AR(1) coefficient across stocks. In other words, it permits some (but not all) of the individual series to have unit roots. Finally, we employ the Maddala and Wu (MW, 1999) non-parametric test, which uses the test statistic $-2\sum_{ij=1}^{N} \ln(\Pi_{ij})$, where Π_{ij} is the p-value from the DF or ADF tests for fish each stock ij.

The LLC and IPS tests require a balanced panel and our panel is unbalanced because some fish stocks entered at different times; for instance, rock lobster enters the quota management system in 1990–91 (see Table 1 for the differences in series length).[23] To address this issue, we conduct the tests for the periods 1990–2001 and 1987–2001, where the former includes most of the shellfish fisheries (but covers a shorter length) and the latter excludes shellfish (but uses a longer panel).

For each of these two panels, we undertake the tests at the fish-stock level and also at the species level (using aggregate species-level price indices).[24] For the LLC and IPS tests, we include one lag, a constant term, and a common (panel-wide) time trend. The IPS test is done on the cross-sectionally demeaned data (extracting panel-specific time effects). The MW test is similar to the LLC, except that the coefficient on the time trend is allowed to vary across the panel.

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[23] The MW test does not require a balanced panel, so we ran the tests for the entire period including all fish stocks regardless of the date at which they entered the system. We also estimated the test statistic on the same sample as the IPS and LLC to compare the results.

[24] The indices are a weighted average of the quarterly fish stock prices, where the weights are equal to each fish stock's share of the total species TAC.

We reject the null hypothesis of a panel unit root for the quarterly asset and lease prices at both the fish stock and species level aggregations in each of the tests at the 1 percent level.[25] The results are the same regardless of the panel length. The agreement in the time series properties of the asset and lease prices satisfies, at least at the panel level, a necessary condition of the present-value model (Falk 1991).[26]

An advantage of the MW test is that it is based on the p-values for each individual series in the panel. Running the test therefore also provides the distribution of p-values from the individual series ADF tests, which we find exhibit substantial heterogeneity at both the species and fish stock levels.[27] While understanding the differences in the time series properties across the markets is an interesting question, it falls beyond the scope of the current paper.

Finally, we also test for the possibility of non-stationarity in the quarterly New Zealand real interest rate for 30-day Treasury bills and the quarterly species-level export price, which is used as an instrument in the econometric analysis for contemporary lease prices. In both cases, we reject the null hypothesis of a unit root. Therefore, the time series variables in the regressions to follow are all stationary, allowing us to draw inferences from the use of standard panel data techniques with variables in levels.

Panel Estimation Techniques

Because lease prices and asset prices are determined simultaneously in the ITQ market

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[25] Each of the above tests allows serial correlation in the errors, but the LLC and IPS are based on the assumption of independence of the errors across the fish stocks. Using Monte Carlo simulations, Bornhorst (2003) shows that depending on the nature of the dependence, either short-run correlations or long-run cointegration relationships, the LLC and IPS tests can lead to both type I and type II errors. It is possible that fishing production relationships among stocks (bycatch, for example) might lead to the violation of the error independence assumption. Given the flexibility in the market and especially the role of deemed value payments (see Footnote 6), it is not clear that these production relationships will be empirically measurable in asset and lease prices during our sample. We did, however, check for possible distortions in our test results by running the same battery of tests after removing the New Zealand Ministry of Fisheries list of Quota Management System species that are caught jointly. We again reject the null hypothesis of a unit root across all tests at the 1 percent level.

[26] A testable implication of the present value model is that the time series properties of asset prices and dividends should be identical. That is, if rents are (non-)stationary, then agricultural land prices should be (non-)stationary. Falk (1991) and Clark *et al*. (1993) test this necessary condition for Iowa, U.S., and Illinois farmland, respectively. Falk (1991) finds that in the Iowa market both series follow a unit root, while Clark *et al*. (1993) reject the hypothesis that the two series have the same time series representations.

[27] For example, using the species-level price indices, we find a median p-value of 0.02 and a mean of 0.11 with a standard deviation of 0.18. The lease prices have a median p-value of 0.01, a mean of 0.079 and standard deviation of 0.156.

each period, it is likely that estimation of equation (6) suffers from simultaneity bias. Statistical tests for endogeneity of the lease price verify this concern.[28] We therefore use instrumental variables estimation throughout, instrumenting for the log lease price using the logged contemporaneous export price of fish and all other regressors, including stock fixed effects. The price of fish is an excellent instrument as it is a significant determinant of profits from fishing, it is highly correlated with quota lease prices ($\rho = 0.77$), and it is exogenous.[29] Our estimation approach follows Balestra and Varadharajan-Krishnakumar (1987), employing a two-stage least squares generalization of the standard panel data estimators to correct for endogeneity. See Baltagi (2001) for an introduction to panel data models with endogenous explanatory variables.

Our first specification models the data for all stocks in a pooled fashion (see Table 3 for results). This approach is appropriate if there are no unobserved stock-specific effects. In contrast, our second specification performs the within estimator, which is equivalent to a regression with a full set of stock-specific fixed effects. While the within estimator is consistent, it is not as efficient as other estimators (*e.g.*, random effects) if the unobserved stock-specific effect is uncorrelated with the observed regressors. In addition, with the fixed effects approach it is not possible to recover coefficients on any of the time-invariant regressors, namely the export price growth rate, mortality rate, and recovering stock dummy.

Our third specification performs the between estimator, which is a regression of stock-specific averages over time. As such, this specification cannot identify coefficients on stock-invariant regressors, such as the interest rate. Finally, our fourth specification treats the stock-specific term, v, as a random effect.[30] This model is more efficient than within estimation when none of the regressors are correlated with the stock-specific effect, however it is inconsistent when the opposite is true. This assumption of no correlation is typically assessed using a Hausman test. The random effects estimator has the advantage of controlling for stock-specific effects while at the same time allowing for estimation of time-invariant explanatory variables,

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[28] Both the Wu-Hausman F test and Durbin-Wu-Hausman chi-squared test strongly rejected (p-value < 0.0001) the null hypothesis of exogeneity of the lease price (see Davidson and MacKinnon 1993). Rejection indicates that an instrumental variables estimator should be employed.

[29] It is reasonable to treat fish prices as given because New Zealand exports about 90 percent of its commercial catch, yet accounts for less than 1 percent of world fishing output. Even in the small number of cases in which New Zealand comprises a sizeable fraction of the world catch of individual species, these species have many near-perfect substitutes in the form of other "white fish."

[30] The random effects estimator is an efficient weighted average of the within and between estimators, assuming its assumptions are upheld.

which are of central interest to this paper. Beyond the fixed or random effect, we assume the remaining error is homoskedastic. This is supported by panel tests of both heteroskedasticity (Pagan and Hall 1983) and autocorrelation (Wooldridge 2002), neither of which rejected a homoskedastic error structure.

Estimation Results

In Table 3, we report the results of estimating equation (6) using the fishing quota asset and lease price data described above. Due to some observations having missing values for one or more of the variables, 4,120 observations were available for the four models that are reported. All specifications show a high degree of explanatory power, with R^2 values at 0.77 or above. Overall, the results are largely consistent with economic expectations about the parameters. The estimated coefficients generally have the expected signs and reasonable magnitudes and are consistent across the different specifications.

Regarding the appropriateness of the different specifications, we find that a joint F-test of the fixed effects is highly significant, thereby opening up the standard errors and consistency of the pooled model to misspecification. On the other hand, a Hausman test comparing the fixed effects and random effects estimators clearly indicated that the random effects model was appropriate (*i.e.*, the assumption of no correlation between regressors and the random effect was not rejected). Hence, the random effects estimates are consistent and more efficient than the fixed effects (or between) estimates. Indeed, the stability of the parameter estimates across these specifications illustrates the consistency of the random effects model.

All four specifications reported in Table 3 feature an estimated coefficient on log lease price of between 0.76 and 0.86. As expected, the random-effects estimate for the lease price lies between the within estimate (model ii) and the between estimate (model iii). These results suggest that changes in lease prices are reflected very closely in changes in the contemporaneous quota asset prices. However, the point estimates are somewhat lower than the expected coefficient of 1 based on the simple present-value model given above, or based on the simple univariate relationship depicted in Figure 2.[31]

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[31] A t-test does not reject the null that the coefficient on lease price is equal to 1 for specification (ii), but it does so for specifications (i), (iii) and (iv).

It is not clear how much this lower-than-expected estimated coefficient casts doubt on the simple present value model represented by equation (6), although it suggests that it may not hold exactly. One possibility is an errors-in-variables problem, with the standard implication that the resulting coefficient is biased toward zero. Another possibility is that all the various controls included in the model (*e.g.*, year effects) are simply reducing the amount of variation in lease prices available for estimating that coefficient. This conjecture is supported by a simple random or fixed effects regression of log quota asset prices on (instrumented) log quota lease prices (with no other controls). In these simple models the coefficient on the log lease price is 0.98 in the case of the random effects model and 1.04 in the fixed effects model, with neither of these estimates being statistically different from 1.

A further possibility is that quota prices adjust slowly in response to changes in profit conditions, and that the contemporaneous lease price is an insufficient indicator of expectations about future profits. This possibility could warrant the inclusion of multiple lagged values of the lease price in the estimation equation, as in the model of Burt (1986). We explored this by including the one-year and two-year lagged lease price in the fixed and random effects models, finding that these lagged prices were statistically insignificant and did not increase the total effect of lease prices on asset prices. In addition, we explored an adaptive expectations model (as described in Section 3) by including lagged asset prices as a regressor and estimating the model according to the approach of Arellano and Bond (1991) to account for the lagged dependent variable. The estimated coefficient on the lagged asset price was 0.08 and was statistically insignificant from zero, again suggesting that using the contemporaneous price in conjunction with the other variables affecting expectations (*e.g.*, export price growth) is acceptable.

In general, the estimated coefficients on the other regressors are consistent with the predictions of the theory outlined in Section 3. Periods with higher interest rates have lower asset prices, *ceteris paribus*, as predicted by the basic present value model. Stocks with more uncertain returns, in the form of higher mortality rates, also tend to have lower asset prices, as would be expected in the presence of risk-averse fishermen.

With respect to the magnitude of these estimates, we refer back to the implications of the strict interpretation of the logarithmic approximation given by equation (6), which are $\beta_2 \approx -(1+r)/(r+\theta-g)$, and $\beta_3 \approx -\theta/(r+\theta-g)$, where each of the variables in these formulae is taken to be its mean value. For $r = 6.4\%$ and $r+\theta-g = 8.9\%$ (based on the mean lease-to-asset price ratio), we would expect $\beta_2 \approx -12$. Our estimate is $\hat{\beta}_2 = -4$, which is in the same realm, but somewhat muted relative to what the simple theory suggests. At the same time, a risk

premium of $\theta = 3.8\%$ (based on $\theta = 8.9\% - r + g$ and $g = 1.3\%$) yields $\beta_3 \approx -0.4$, which is similar to our estimate of $\hat{\beta}_3 = -0.3$. Note that although we do not have a direct measure of the risk premium, the mortality rate proxy we use should yield approximately the same estimated coefficient if it is directly proportional to the true measure of lnθ.

Table 3 also reports evidence that stocks with faster-growing returns have higher asset prices, controlling for other factors. As noted above, growth in returns may be due to rising prices or falling costs. The former clearly has an important impact on New Zealand fishing quota asset prices, as growth in export prices is found to be strongly associated with asset prices in all specifications where this effect could be estimated. Regarding the magnitude of this effect, earlier we set out the hypothesis that $\beta_4 \approx (1 + g)/(r + \theta - g)$, which yields $\beta_4 \approx 12$, while our estimate is approximately $\hat{\beta}_4 = 4$. Interestingly, while the estimated coefficients on $\ln(1 + r)$ and $\ln(1 + g)$ are, as the theory suggests, approximately equal and opposite in sign, both are lower than the theoretical expectation. One possible explanation is measurement error (errors in variables problem) leading to the usual bias towards zero.

Stocks where fishing costs are expected to fall over time are also found to have higher asset prices. This is seen in Table 3 in two ways. First, recovering stocks tend to have higher asset prices, as expected. Contrary to our hypothesis, however, we found no evidence that this premium has dissipated over time, with a very small and statistically insignificant coefficient on the time trend found for recovering stocks. One explanation for this finding may simply be that the expected future recovery of these stocks has yet to be fully realized, due to the life-cycle characteristics of the fish populations and ocean environmental conditions.

Second, we found that high levels of trade in the quota asset market were associated with higher asset prices across stocks, after controlling for other effects, but that this effect was statistically insignificant. The positive point estimate is consistent with the notion that stocks experiencing a high degree of rationalization after the introduction of the quota system feature decreasing fishing cost and thus become increasingly valuable over time. Finally, the shellfish dummy (*i.e.*, for rock lobster, abalone, and scallops) is found to enter specifications (i), (iii) and (iv) with a highly significant positive coefficient. This suggests that shellfish stocks tend to have higher asset prices than other stocks, *ceteris paribus*. One possible explanation for this additional effect of shellfish stocks is that the biomass of these species is typically estimated with more

precision, and, hence, their catch rates are less uncertain.[32]

Conclusion

Individual fishing quotas are a promising market-based system for avoiding the common pool problem in fisheries, particularly when trade of quotas between fishers is permitted. When there are competitive quota markets, rational asset pricing theory suggests that the price of quotas should reflect the expected present value of future profits in the fishery. Thus, for ITQs to deliver an efficient solution to the common pool problem, quota markets must convey appropriate price signals.

The aim of this paper is to extend the analysis of Newell *et al*. (2005) by econometrically examining the relationship between the quota asset and annual lease prices in the New Zealand ITQ market. Random effects and other panel models revealed that quota asset prices are related in the expected manner to contemporaneous lease prices. We also found that asset prices are higher when interest rates are low and for stocks that experience less biological fluctuation. Furthermore, stocks with higher growth rates of fish output prices tend to have higher quota asset prices. Stocks that are thought to have experienced reductions in costs since the introduction of the ITQ market are also found to have higher asset prices, *ceteris paribus*. However, these effects were not found to have decreased over time as expected.

By implication, the New Zealand quota system as a whole, has functioned reasonably well and the prices at which quotas have sold appear to reflect expectations about future returns on specific fish stocks.

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[32] There is also anecdotal evidence that these stocks tend to have more effective cooperative management institutions (*e.g*., Challenger Scallop Enhancement Company), which is likely due to the biological and production characteristics of the stocks.

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Tables and Figures

Table 1. Species included in the New Zealand ITQ system as of 1998

Species	Abbreviation	Year entered	Fish stocks	Species type
Barracouta	BAR	1986	4	Offshore
Blue cod	BCO	1986	7	Inshore
Bluenose	BNS	1986	5	Inshore
Alfonsino	BYX	1986	5	Inshore
Rock lobster	CRA	1990	9	Shellfish
Elephant fish	ELE	1986	5	Inshore
Flatfish	FLA	1986	4	Inshore
Grey mullet	GMU	1986	4	Inshore
Red gurnard	GUR	1986	5	Inshore
Hake	HAK	1986	3	Offshore
Hoki	HOK	1986	1	Offshore
Hapuku and bass	HPB	1986	7	Inshore
John Dory	JDO	1986	4	Inshore
Jack mackerel	JMA	1987	3	Offshore
Ling	LIN	1986	7	Offshore
Blue moki	MOK	1986	4	Inshore
Oreo	OEO	1986	4	Offshore
Orange roughy	ORH	1986	7	Offshore
Oyster	OYS	1996	2	Shellfish
Paua (abalone)	PAU	1987	10	Shellfish
Packhorse rock lobster	PHC	1990	1	Shellfish
Red cod	RCO	1986	4	Inshore
Scallops	SCA	1992	2	Shellfish
School shark	SCH	1986	7	Inshore
Gemfish	SKI	1986	4	Offshore
Snapper	SNA	1986	5	Inshore

Rig	SPO	1986	5	Inshore
Squid	SQU	1987	3	Offshore
Stargazer	STA	1986	7	Inshore
Silver warehou	SWA	1986	3	Offshore
Tarakihi	TAR	1986	7	Inshore
Trevally	TRE	1986	4	Inshore
Blue warehou	WAR	1986	5	Offshore

Table 2. Descriptive statistics for determinants of fishing quota asset prices

Variable	*Mean*	*Std. dev.*	*Min.*	*Max.*
Lease price ($/ton)	1,795	4,289	1	43,663
Asset price ($/ton)	20,266	46,870	22	358,586
Export price ($/ton)	8,319	12,096	630	61,009
Export price growth rate	0.013	0.023	-0.027	0.071
Interest rate	0.064	0.022	0.027	0.110
Normalized percentage of quotas sold	1.000	0.952	0.000	11.892
Natural mortality rate	0.222	0.174	0.045	1.000
Reduced TAC (dummy indicating fishery had initial reductions)	0.273	0.446	0	1
Shellfish (dummy indicating shellfish quota market)	0.116	0.320	0	1
Number of leases per quarter	17	20	1	194
Number of asset sales per quarter	4	4	1	75

Note: Statistics are based on the 4,120 observation sample from the estimation of quota asset price determinants. Monetary figures are year 2000 New Zealand dollars, which are typically worth about half a U.S. dollar. Tons are metric tons.

Table 3. Determinants of fishing quota asset prices

Variables	(i) Pooled	(ii) Within (fixed effects)	(iii) Between	(iv) Random effects
Logged lease price (instrumented) (π)	0.840*** (0.016)	0.861*** (0.234)	0.764*** (0.046)	0.833*** (0.036)
Interest rate ($\ln(1+r)$)	-3.048* (1.871)	-3.966** (1.961)		-3.707** (1.746)
Natural mortality rate ($\ln\theta$)	-0.331*** (0.072)		-0.382* (0.209)	-0.333** (0.167)
Growth in output prices ($\ln(1+g)$)	2.927*** (0.593)		4.029** (1.681)	3.725*** (1.358)
Normalized percentage of quotas sold (s)	0.013 (0.012)	0.013 (0.013)		0.012 (0.011)
Fisheries with initial reductions in TAC (a) (dummy variable)	0.099* (0.054)		0.192** (0.077)	0.124* (0.074)
Interaction of time with variable indicating fisheries with initial reductions in TAC ($a \bullet t$)	0.006 (0.006)	0.006 (0.008)		0.007 (0.006)
Shellfish (dummy variable)	0.272*** (0.056)		0.595*** (0.155)	0.319*** (0.123)
Seasonal effects	*jointly significant*	*jointly significant*	—	*jointly significant*
Fish stock fixed effects	—	*jointly significant*	—	—
Year effects	*jointly significant*	*jointly significant*	—	*jointly significant*
Constant term	3.591*** (0.205)	3.567*** (1.320)	3.855*** (0.285)	3.646*** (0.266)
Number of quarterly observations	4,120	4,120	4,120	4,120
Number of panels (quota markets)	152	152	152	152
R-squared	0.78	0.77	0.78	0.79

Note: The dependent variable is the logged average quarterly asset price. Standard errors are shown in parentheses. ***indicates significance at the 1 percent level, ** at the 5 percent level, and * at the 10 percent level. The data are a panel of observations for species- and region-differentiated quota markets over 15 years. Estimation method is indicated above each column.
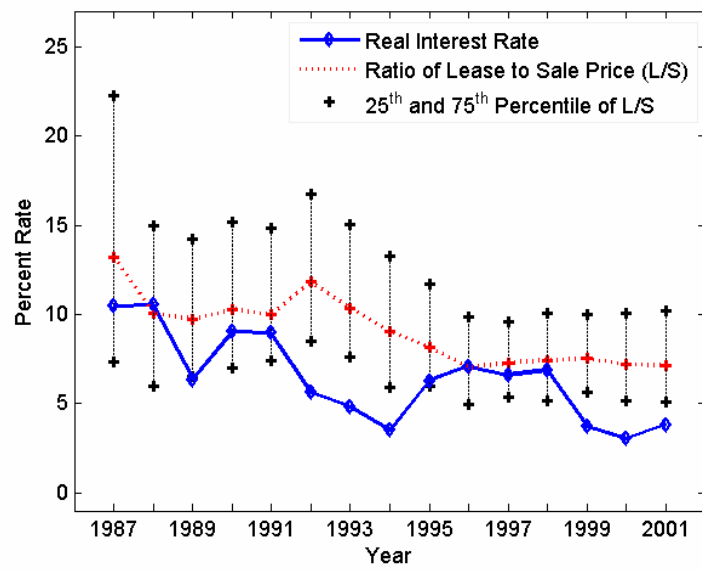
Figure 1. Quota dividend-price ratio and market interest rates

Note: Rates of return are medians (50th percentile), 25th percentile and 75th percentiles across fish stocks in each year. The real interest rate is based on New Zealand Treasury bills, deflated using the New Zealand consumer price index.
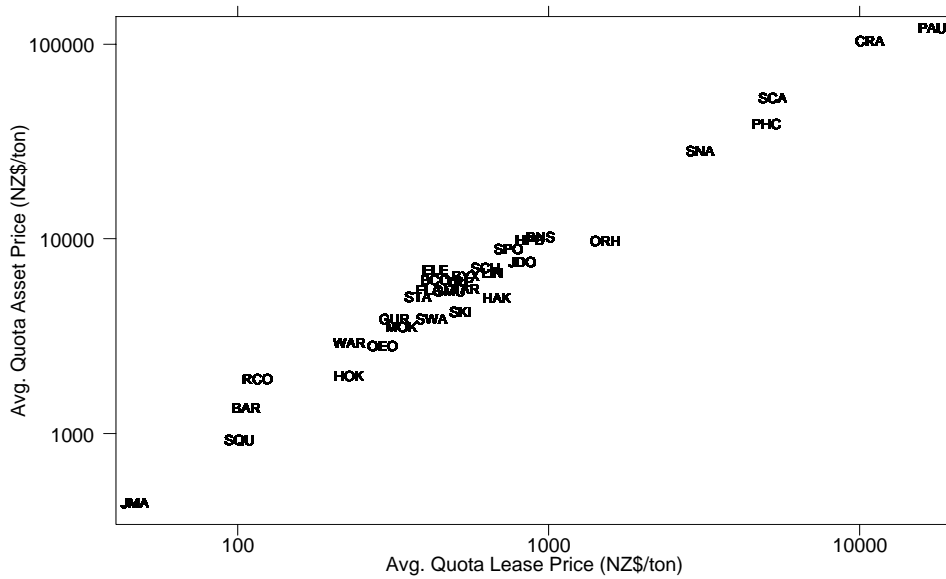
Figure 2. Average quota asset price versus average quota lease price

Note: Logarithmic scale. Averages by species. Year 2000 NZ$. Data symbols are species abbreviations (see Table 1). Note that the asset price and lease price are approximately linearly related with a slope of 1. The level of the asset price is also approximately 10 times the lease price, roughly equal to the present value of a perpetuity discounted at 10 percent.