Simulated Maximum Likelihood Estimation and Analysis of Covariance in a Panel Tobit Model of California's Groundfish Trawl Fishery, 1981-2001

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Michael Dalton

Divison of Science and Environmental Policy

California State University, Monterey Bay

100 Campus Center, Seaside CA, 93933

Email: Michael_Dalton@csumb.edu

Abstract

Spatial management is currently an important issue in fisheries, and a central question for managers is how fishing effort will respond to marine reserves and other types of closures. This paper develops a panel Tobit model to analyze the influence of spatial and dynamic factors on decisions about where and when to fish. The model includes autocorrelation. A simulated maximum likelihood approach is used to compute parameter estimates and conduct hypothesis tests, including an analysis of covariance to detect sources of individual heterogeneity. The model is used with ten panels of data, representing fleets from ports in California's groundfish trawl fishery. Results show that ex-vessel prices are the most important explanatory variable in the model, and affect the spatial distribution of fishing effort. Regulatory variables, in the form of limits on landings for some species, are also important in most cases, and these reveal both spatial and temporal effects of past regulations. Dynamic factors such as autocorrelation, or effects of past fishing effort in a particular area on current effort, are also significant at several ports, but spatial interactions in effort are important in only two cases. Results from the analysis of covariance show that using pooled time series data to analyze effects of spatial management is acceptable practice in some cases.

1. Introduction

Fisheries are an important source of income for many coastal communities. Starting in 1999, the National Marine Fisheries Service declared several West Coast rockfish species to be overfished under standards of the Sustainable Fisheries Act (SFA). In response, the Pacific Fishery Management Council (PFMC) implemented stringent regulations to prevent further overfishing and rebuild rockfish stocks, with severe economic consequences for many communities. In 2001, the entire West Coast groundfish fishery was declared a Federal Disaster. Managing this fishery is a complex task. The fishery's management plan regulates commercial and recreational fleets, includes several types of gear, and covers more than ninety species in three states. Bottom trawling is among the most important sectors of the fishery, and is managed closely through a limited entry program. Historically, the PFMC used bimonthly limits on landings to regulate trawlers. The economic objective of these limits was to produce a uniform flow of groundfish over the year to meet total allowable catch for each stock. However, bycatch is a problem for trawlers, and as limits on rockfish landings decreased, regulatory discard by these vessels became critical. To avoid closing the fishery, the PFMC implemented a Rockfish Conservation Area (RCA) that closed large parts of the continental shelf to bottom trawling and other types of fishing. The RCA was established to eliminate, or drastically reduce, levels of fishing effort in areas with the highest levels of rockfish bycatch.

The RCA is expected to remain intact during the rebuilding period for rockfish stocks, which could last several decades. During this period, boundaries of the RCA may be adjusted, depending on stock status and other factors. In some areas, permanent closures to all types of fishing, known as marine reserves, are being proposed. An important question for fishery managers in these cases is how the spatial distribution of fishing effort would respond to the proposed changes. To address this question, this paper develops an econometric model for individual vessels that is both dynamic and spatial, which is estimated and tested using panel data for groundfish trawlers in California. The model in this paper is new because it treats fishing effort as a continuous decision variable in a regression framework that considers fishing effort to be a censored dynamic variable using a panel Tobit model.

Uncertainty is a fundamental part of many fisheries that can affect decisions about fishing effort. Marine economics contains a substantial literature on effects of uncertainty, and this topic is also important in other areas of economics. The task for the model in this paper, however, is rather narrow: To test for effects of individual heterogeneity in panel data using a dynamic linear model of economic behavior that is motivated explicitly by its relationship to a class of linear rational expectations models. Rational expectations models are attractive for econometric policy analysis because they estimate structural parameters, and thus, attempt to avoid the Lucas Critique that parameter values would change in response to a proposed policy. The linear class of models has been used in many contexts to link macroeconomic outcomes to the microeconomic decisions that generated them. Sargent (1978, footnote 4) refers to this convenient property of linear rational expectations models as admitting a tidy theory of aggregation.

In a fisheries setting, some studies explicitly model how information available to fishermen may be used to make forecasts of future variables that affect current decisions (Rosenman, 1986; Rosenman, 1987; Rosenman and Whiteman, 1987; Dalton, 2001; Dalton and Ralston, 2004). Rosenman (1986) analyzed theoretical implications of the rational expectations hypothesis for fisheries management in a bioeconomic fisheries model. Rosenman (1987) applied the bioeconomic model to fisheries data using rational expectations to derive a set of identifying restrictions. Dalton (2001) extended this type of analysis to test the rational expectations hypothesis. Recently, Dalton and Ralston (2004) developed a spatial version of the rational expectations model, and applied it to pooled time series data for groundfish trawlers to analyze effects of the RCA on trawlers operating at a single port in California. Bioeconomic models under rational expectations are tractable, but sources of bias may exist in the pooled data that have been used for estimation and testing (Hsiao, 1986).

Several recent studies use discrete choice random utility models (RUMs) to analyze spatial distributions of fishing effort from panel data that incorporate effects for individual fishermen (e.g. Holland, et al., 2004). These models provide an appropriate and tractable framework for modeling spatial management and other issues in many fisheries. However an alternative approach, based on time series methods, may be better suited in some cases for analyzing fishery dynamics. For example, open access is sometimes used to justify an assumption that current decisions do not depend on expectations about future conditions, thus profit maximization for individuals is essentially a static decision in some models. While the assumption of open access is plausible in many fisheries, groundfish trawlers on the West Coast are part of a limited entry program, and ignoring information about future conditions for regulations, stock abundance, or climate would not be optimal. In addition, Rosenman (1986) showed that a type of open access equilibrium can occur with behavior that is forward looking, and the dynamic policy implications for fishery managers in this case are different from those of a static model. Therefore, assumptions both about dynamic behavior and the presence of individual heterogeneity should be tested. Practical experience supports this type of testing: Fishermen on the West Coast are known to modify behavior based on expectations of future conditions. Therefore, forward looking behavior is a plausible response to uncertainty about future regulations, price changes, climate fluctuations, or other events.

The dynamic and spatial model of fishing effort in this paper allows for heterogeneity at the level of individual fishermen. The model is used for estimation, and for an analysis of covariance to test parameter restrictions across individuals in panel data. The model is based on a pair of dynamic equations, and is not yet ready to test the cross equation parameter restrictions implied by rational expectations. Instead, this paper addresses the econometric issues that arise in applying these equations to panel data. The rational expectations models cited above give a structural relationship between fishing effort, the endogenous variable in the dynamic equations used in this paper, and expected profits of individual fishermen. In technical terms, solutions to the dynamic optimization problem in these models are interior solutions to the stochastic Euler equations. Therefore, estimation and testing has been restricted to data with positive levels of fishing effort for each year and location in the analysis. While past work has used pooled data to satisfy this requirement, the conditions necessary for obtaining unbiased estimates from these data have not been tested. However, zero values in a particular year or location is a common event for individual vessels. The econometric issue in these cases is that error terms are correlated with the regressors, and thus, estimates may be biased.

An event with zero fishing effort implies that fishing occurred at another location, or that no fishing occurred, either because a different activity yielded a higher expected return, or because of a random factor such as weather or vessel maintenance. In a dynamic optimization framework under rational expectations, decisions about whether to fish, and if so, how much effort to use could be expressed as a set of inequalities or Kuhn-Tucker conditions. The standard Euler equations are a special case of equality for these conditions that imply interior solutions. Zero effort would be optimal only if the corresponding Kuhn-Tucker condition were a strict inequality. In this case, expected profits are not observed. Therefore, fishing effort can be treated as a censored variable, and a dynamic Tobit model used for estimation and testing.

Maddala (1983) shows a standard maximum likelihood procedure for the Tobit model gives unbiased estimates for a fixed-effects dynamic model with censored endogenous variables, observed lagged variables as regressors, and independent disturbances. However, Lee and Maddala (1985) show that even first order autocorrelation in the disturbances with censored lagged variables as regressors can create severe computational problems for the maximum likelihood approach. The maximum likelihood procedure in this case involves computation of integrals up to the same order as the number of censored variables in the data set. The literature on discrete-choice models deals with similar computational problems because likelihood functions for these models involve integrals of the same order as the number of time periods in the data set. Lee (1998) demonstrates these computational problems can be overcome with a simulated maximum likelihood approach by developing a version of the Geweke-Hajivassiliou-Keane (GHK) likelihood simulator (Börsch-Supan and Hajivassiliou, 1993; Keane, 1994; Geweke, 1991) for Probit models that use panel data for estimation, and evaluates performance of the likelihood simulator using Monte Carlo methods. Lee (1999) continues this assessment of the simulated maximum likelihood approach for models with limited dependent variables by developing another version of the GHK simulator for dynamic Tobit models.

The panel Tobit model presented in this paper is an extension of the dynamic Tobit model in Lee (1999), for use with panel data, as in Lee (1998). A version of the GHK likelihood simulator is also presented for the panel Tobit model. Monte Carlo tests have been performed using a likelihood simulator similar to the one presented in this paper (Dalton, 2005). Results from these tests show maximum likelihood estimates are most accurate with panels of at least forty individuals and sixty time-periods. However, the Monte Carlo results also show results of the analysis of covariance are reliable in panels with only about ten individuals and twenty time-periods, which is satisfied in most cases by the fisheries data used for analysis in this paper. The panel Tobit model used in this paper is spatial, but limited to just two areas: Inside the RCA, and the area outside that is not subject to other closures. For example, trawling within two miles of the California coast is prohibited. A more refined analysis is possible in future work, but even the two-area model presents computational challenges, and it provides a useful starting point.

Data for groundfish trawlers at ten California ports are used with the panel Tobit model to analyze dynamics of fishing effort with spatial movement. The scope of analysis is limited to a common fishing strategy used by groundfish trawlers in California, which is to target the dover sole-thornyheads-sablefish (DTS) complex. Data for the ten ports are treated as separate panels with fishing effort and ex-vessel prices for individual trawlers in the two areas over twenty-one years, 1981-2001. More recent data have been collected, but important structural changes occurred after the 2001 fishery, including establishment of the RCA, and more recently, a trawl buyback program that reduced the size of the fleet. Future work will compare model results through 2001 to more recent data, but because of confidentiality restrictions, access to these data is tightly controlled. In addition, working out changes in vessel ownership and limited entry permits from the trawl buyback program will be a detailed and data-intensive task. Data for each port is associated with an area that gives the total spatial extent of trawl effort in the panel.

A geographical information system (GIS) was developed that includes bathymetric (depth) data for the U.S. Exclusive Economic Zone (EEZ), and the spatial location of each tow from information in trawl logbooks. The bathymetric data and tow locations were linked, and queried to identify tows for individual vessels that occurred inside and outside of the original RCA that was established by the PFMC after the 2001 fishery. Another source of data in PacFIN, fish ticket receipts from landings, was used to construct time series of ex-vessel prices for each vessel. Ex-vessel prices are derived from fish tickets for each port as the ratio in each year of total DTS revenues divided by total DTS landings. In general, due to differences in size and quality of fish, a price differential may exist for fish caught in different areas but this differential is not included in the analysis because the quality of spatial data for constructing ex-vessel prices in California is questionable, and further work with these data is needed to evaluate this issue.

Results of estimation and analysis of covariance with the panel Tobit model are reported in the paper for each of the ten ports. With twenty-one years of data and fewer than forty individuals at any port, the model still yields reasonable estimates for both individual and pooled versions of the model. For example, coefficients on ex-vessel prices are positive, and typically significant. Point-estimates for coefficients on lagged effort variables have absolute values less than one, which are necessary conditions for model stability, and are significant in many cases. Sufficient conditions for stability, which essentially bound the parameters that describe interactions between areas, are also met. Autocorrelation in the error terms is present, and significant in several cases.

For the analysis of covariance, hypothesis tests compare likelihood values from individual and pooled versions of the panel Tobit model. The individual version of the model allows each vessel to have a separate coefficient for each parameter. In the pooled version of the model, each parameter is restricted to have the same value for all individuals. The set of parameters for each vessel includes coefficients on ex-vessel prices, lagged effort inside the RCA, and lagged effort outside the RCA. Other parameters describe variance and autocorrelation. Restrictions imposed by the pooled model are rejected at the 5% level for four of the ten ports that were analyzed. For these four, an additional likelihood ratio test was conducted using a cell corrected model that estimates at the individual level a constant and the coefficient on ex-vessel prices, which is the most important explanatory variable in the model. This cell corrected model is not rejected at these four ports. A similar set of tests was conducted using a model with individual heterogeneity only in a constant term, but this model has convergence problems, presumably because this term is not significant in most cases.

A second set of results with the panel Tobit model includes effects of regulations in the form of trip limits for widow rockfish as an additional exogenous variable. Trip limits for several species, including sablefish and others in the DTS complex, were analyzed and the series for widow rockfish performed best. Some results described above are sensitive to including regulatory variables. The pooled version of the policy model is rejected for six ports, and a cell corrected version of the policy model that allows individual heterogeneity in each of the exogenous variables is rejected at the 5% level for two ports. Therefore, an analysis of covariance cannot rule out significant individual effects in the dynamic parameters of the policy version of the panel Tobit model. The policy version of the model yields reasonable estimates for coefficients on ex-vessel prices and the model's dynamic parameters. Effects of the rockfish regulations are significant at eight of the ten ports in the analysis, and except for one port, maximum likelihood estimates of the coefficient on the regulatory variable are positive, implying that reducing trip limits for widow rockfish decreases fishing effort for DTS species.

Sec. 2 of the paper presents the panel Tobit model and likelihood simulator. Sec. 3 describes the fisheries data that are used for estimation and testing. Sec. 4 gives results for maximum likelihood estimates and the analysis of covariance at each of the ten ports included in the analysis using the model with ex-vessel prices as the only exogenous variable. Sec. 5 gives these results for the policy version of the model that includes ex-vessel prices and regulations as exogenous variables. Sec. 6 discusses the panel Tobit model and results, and provides some directions for future research.

2. Panel Tobit Model and Likelihood Simulator

Observed endogenous variables for individuals $i = 1, \ldots, N$ in periods $t = 1, \ldots, T$, and in areas $k = 1, 2$ are given by $y_{kit} \geq 0$ for each k, i, and t. The underlying latent variable before censoring is y^*_{kit}, and the relationship with the observed dependent variable is given by $y_{kit} = \max\{y^*_{kit}, 0\}$. Exogenous variables for individuals are given by the collection x_{it}. Let ω_{kit} denote a collection of normal random variables, each with mean zero and variance σki^2, such that $\epsilon_{kit} = \rho_{ki}\epsilon_{kit-1} + \omega_{kit}$ for all k, i, and t. Assume the process is invertible with $-1 < \rho_{ki} < 1$. Given the process for ϵ_{kit}, and other conditions described above, a first-order linear panel model is described by coefficients α_{ki}, β_{ki}, λ_{ki}, and γ_{ki}, parameters ρ_{ki} and σ_{ki}, and a pair of regression equations for each i,

$$
\begin{aligned}
y^*_{1it} &= \alpha_{1i} + \beta_{1i}x_{it} + \lambda_{1i}y^*_{1it-1} + \gamma_{1i}y^*_{2it-1} + \epsilon_{1it}, &(1)\\
y^*_{2it} &= \alpha_{2i} + \beta_{2i}x_{it} + \lambda_{2i}y^*_{2it-1} + \gamma_{2i}y^*_{1it-1} + \epsilon_{2it},\ t = 1, \ldots, T. &(2)
\end{aligned}
$$

The fixed effects model above is a member of the general framework in Lee (1999), which assumes that $y^*_{kit} = \Psi(\bar{y}_{kit-1}, \bar{y}^*_{kit}) + \omega_{kit}$, for a function Ψ, where $\bar{y}_{kit}$ denotes a vector consisting of the current observed dependent variable y_{kit}, and lagged values up to t, and $\bar{y}^*_{kit}$ denotes a vector consisting of the latent-dependent variables up to t. Define the lag operator L, note that $\epsilon_{kit} = (1 - \rho L)^{-1}\omega_{kit}$, and substitute into (1) or (2) for each $k = 1, 2$.

Let I_{it} denote the indicator function such that $I_{kit} = 1$ if $y_{kit} > 0$, and $I_{kit} = 0$ if $y_{kit} = 0$. Given sample observations, suppose that $I_{kis} = 0$ for all $s \in \{t_1, t_2, \ldots, t_m\} \subseteq \{1, \ldots, T\}$, such that $t_1 < t_2 < \ldots < t_m$, and $I_{kis} = 1$ for s not in $\{t_1, t_2, \ldots, t_m\}$. Then, $\bar{y}^*_{kiT} = (y^*_{kit_1}, y^*_{kit_2}, \ldots, y^*_{kit_m})$, and $\bar{y}_{kiT}$ consists of all remaining uncensored or observed positive values of fishing effort. Let $\bar{I}_T = (I_1, \ldots, I_T)$. The joint density of observable and latent-dependent variables for the model is defined by $f(\bar{y}_{kiT}, \bar{I}_T, \bar{y}^*_{kiT})$ for each $k = 1, 2$. Lee (1999) describes two approaches for simulating the dynamic Tobit likelihood function, and concludes the approach that allows sample feedback in simulation is more efficient. A simulator for the more efficient approach is derived from the factorization in Lee (1999) equation 3.4,

$$
\begin{aligned}
&f(\bar{y}_{1iT}, \bar{I}_{1T}, \bar{y}^*_{1iT}, \bar{y}_{2iT}, \bar{I}_{2iT}, \bar{y}^*_{2iT}) = \\
&\qquad \prod_{k=1}^{2}\prod_{s=1}^{T}\left(f(y_{kis}|\bar{y}_{kis-1}, \bar{y}^*_{kis-1})\right)^{I_{kis}}\left(\Pr(I_{kis} = 0|\bar{y}_{kis-1}, \bar{y}^*_{kis-1})\right)^{1-I_{kis}} \\
&\times \prod_{s\in\{t_1,t_2,\ldots,t_m\}} g(y^*_{kis}|I_{kis} = 0, \bar{y}_{kis-1}, \bar{y}^*_{kis-1}). \qquad (3)
\end{aligned}
$$

This decomposition shows that each $y^*_{kit_1}, y^*_{kit_2}, \ldots, y^*_{kit_m}$ can be drawn recursively from the conditional univariate densities $g(y^*_{kis}|I_{kis} = 0, \bar{y}_{kis-1}, \bar{y}^*_{kis-1})$, which include past as well as current sample information. Lee describes this approach as allowing current sample feedback, which improves the efficiency of the sampler for Tobit models relative to samplers that do not include such feedback. For a finite number of simulation runs R, the sampler g provides an unbiased likelihood simulator

$$\hat{L}(\bar{y}_{1iT}, \bar{I}_{1iT}, \bar{y}_{2iT}, \bar{I}_{2iT}) = \frac{1}{R}\sum_{r=1}^{R}\prod_{k=1}^{2}\prod_{s=1}^{T}\left(f(y_{kis}|\bar{y}_{kis-1}, \bar{y}^{*(r)}_{kis-1})\right)^{I_{kis}}\left(\Pr(I_{kis} = 0|\bar{y}_{kis-1}, \bar{y}^{*(r)}_{kis-1})\right)^{1-I_{kis}}. \quad (4)$$

Lee shows this simulator for the Tobit model is a generalization of the GHK simulator for multivariate normal choice probabilities in discrete choice models (Geweke, 1991; Borsch-Supan and Hajivassiliou, 1993; Keane, 1994).

Combining simulated maximum likelihood approaches for panel data with the dynamic Tobit model for individuals involves the log of the likelihood function in (4), with a sum over individuals

$$\tilde{\mathcal{L}}(\bar{y}_{1iT}, \bar{I}_{1iT}, \bar{y}_{2iT}, \bar{I}_{2iT}) = \sum_{i=1}^{N}\ln\left(\frac{1}{R}\sum_{r=1}^{R}\prod_{k=1}^{2}\prod_{t=1}^{T}\left(f(y_{kis}|\bar{y}_{kis-1}, \bar{y}^{*(r)}_{kis-1})\right)^{I_{kis}}\left(\Pr(I_{kis} = 0|\bar{y}_{kis-1}, \bar{y}^{*(r)}_{kis-1})\right)^{1-I_{kis}}\right). \quad (5)$$

Since the ω_{kit} are normal random variables, terms in (5) are derived from the normal density and distribution functions. The key feature of likelihood simulator with sample feedback is to simulate only latent variables for censored dates $t_1, t_2, \ldots, t_m$, conditional on observed data for the other dates. The Markov structure is convenient for simulating effects of autocorrelation by computing probabilities of rectangles with the multivariate normal distribution, which has been the subject of much numerical work and well-tested routines are available (Geweke, 1991; Hajivassiliou, McFadden and Ruud, 1996).

The log-likelihood function in (5) is a function of α_{ki}, β_{ki}, λ_{ki}, ρ_{ki}, and σ_{ki} for each k and i. A pooled regression requires that $\alpha_{ki} = \alpha_k$, $\beta_{ki} = \beta_k$, $\lambda_{ki} = \lambda_k$, $\gamma_{ki} = \gamma_k$, $\rho_{ki} = \rho_k$, $\sigma_{ki} = \sigma_k$, for each k and i. The cell-corrected model allows separate α_{ki} and β_{ki} for each k and i, but restricts other parameters to values from a pooled regression. The simulation procedure for the model follows Lee (1999), and the algorithm runs as follows: Assume that $t-1$ variables have been simulated, if $y_{kit} > 0$ then $\epsilon_{kit} = y_{kit} - (\alpha_{ki} + \beta_{ki}x_{kit} + \lambda y^*_{kit-1} + \gamma_{ki}y^*_{kit-1})$. In the uncensored case at t, the first term in the product in (5) is calculated from the normal density with variance σ^2_{ki}, $\phi(y_{kit} - (\alpha_{ki} + \beta_{ki}x_{kit} + \lambda_{ki}y^*_{kit-1} + \rho_{ki}\epsilon_{kit-1}))$. Otherwise, $y_{kit} = 0$, and the algorithm proceeds by drawing a random normal deviate ω_{kit} to simulate $\epsilon_{kit} = \rho_{ki}\epsilon_{kit-1} + \omega_{kit}$, and y^*_{kit} are generated from these according to (1) and (2).In the case of censoring at t, the second term in the product in (5) is given by the normal distribution $\Phi(-(\alpha_{ki} + \beta_{ki}x_{kit} + \lambda_{ki}y^*_{kit-1} + \rho_{ki}\epsilon_{kit-1}))$.

The pooled maximum likelihood estimates are compared to results from maximizing a log-likelihood function that allows heterogeneity. Likelihood ratio tests are used to compare restricted and less restricted models, and are based on an asymptotic chi-squared distribution with degrees of freedom equal to the number of additional parameter constraints imposed by the more restricted version of the model (Amemiya, 1985; Hamilton, 1994). Fortran 77 code was developed to implement the estimation and testing procedures. Routines from Press et al. (1996) are used to maximize likelihoods with Powell's method, evaluation of the univariate normal distribution function with Gaussian quadrature, and evaluation of the chi-squared distribution function with the incomplete gamma function. Uniform and normal pseudo random number generators, and inverse normal distribution function, use routines from Hajivassiliou, McFadden and Ruud (1996).

3. Fisheries Data

Data from the Pacific Fisheries Information Network (PacFIN) are used for estimation and testing of the panel Tobit model. PacFIN is an ongoing census, starting in 1981, of commercial fishing vessels that operate off the coasts of California, Oregon, and Washington. Logbook data from the West Coast Limited Entry Groundfish Trawl Fishery has information on species composition and location of individual tows by each vessel in the fishery. Data from the logbooks are linked to PacFIN fish tickets, which are available for all commercial landings in California, Oregon, and Washington, to obtain ex-vessel values and landed weights of different species for each fishing trip. These variables are used to calculate ex-vessel prices.

An untested assumption used in this paper is that ex-vessel prices are the same across vessels at each port in a given year, which may not be accurate at a finer spatial or temporal scale. For example, size or other characteristics of target species may be influenced by location, but spatial information in the fish tickets does not correspond to the logbook records in many cases, which are more reliable because these are registered to vessels, while buyers submit fish tickets. Otherwise, standard economic assumptions imply price-taking behavior on the part of individual vessel operators and prices should be the same across vessels at each port. Testing these assumptions is left for future work. The working assumption in this paper is an ex-vessel price index formed by the total ex-vessel value divided by the total landed weight of DTS species in a year at each port adequately and exogenously describes the return for all vessels there.

Fig. 1 shows ten ports in California's groundfish trawl fishery that are included in the analysis described below, all of which are located north of Point Conception, the southern limit for most trawlers. Note the geographic ordering of ports from north to south because it is used as a format for presenting data and results below. The two curves along the coast in Fig. 1 identify the depth contours that represent the inner and outer boundaries of the RCA in California. The contours in the figure are at 100 and 450 meters, respectively, and are based on bathymetry data available in GIS files from California Department of Fish and Game (CDFG). These depth contours are closest in the bathymetry data to the original depth contours in the RCA of 50-250 fathoms. CDFG collects logbooks, and records data from these for PacFIN. CDFG manages data on the location of individual tows using a system of statistical fishing blocks, composed typically of 10-degree grid cells. GIS files for the statistical fishing blocks are also available from CDFG. The fishing blocks highlighted in Fig. 1 are contained in, or intersect, the depth contours used to represent boundaries of the RCA. These blocks are used to sort data on fishing effort and catch in the trawl logbooks into areas inside and outside of the RCA. Because the 10-degree blocks are coarse relative to bathymetry of the continental shelf, blocks were selected to cover almost the entire area between the inner and outer depth contours in Fig. 1. However, only a fraction of some blocks is in the depth zone of the RCA.

Work in this paper assumes that logbook data were produced by tows inside the RCA if the corresponding fishing block where the tow occurred has more than 10% of its area in the depth zone of the RCA. The rationale for this assumption is that trawling for DTS species is often done by locating a particular depth, starting a tow, and maintaining the starting depth for the duration of the tow. Blocks that intersect the inner boundary are typically close to land, and therefore, also affected by California's ban on trawling within two miles of the coast. Areas outside of the RCA are further from port and less safe. In addition, blocks that intersect the outer boundary of the RCA often have sharp changes in bathymetry along features unsuitable for bottom trawling such as Monterey Canyon, or the edge of the continental shelf, which are both visible in Fig. 1. Therefore, a reasonable assumption for fishing blocks that intersect the depth contours in Fig. 1 is that trawling was concentrated in the area now inside the RCA.

The data used for the analysis in this paper consist of annual time series for ex-vessel prices and fishing effort at each port, inside and outside of the RCA, for individual trawlers on trips that yielded DTS landings at the ports in Fig. 1. Aggregating to annual levels is for convenience, and avoids effects of seasonality in the underlying data. Logbook records provide information on species composition and duration of individual tows for each fishing trip by a vessel. The total duration over a year of tows that yielded DTS species within an area for an individual vessel on trips that landed at a particular port is the basic unit of data for analysis in this paper. These data form a time-series of ex-vessel prices for each port that is associated with a panel of fishing effort, measured in terms of tow hours per year, inside and outside of the RCA.

The maximum number of individuals at a single port is thirty-seven, and the minimum is four. The condition for associating a vessel with a port are records for at least six years of landings there, which is generally necessary for convergence of the estimation procedure. A separate analysis is conducted for each port. Effort levels at different ports by a single vessel are assumed to be independent events, but the effects of this assumption are probably small. Of the one hundred vessels considered in the panel, seventeen satisfy the landings condition at two or more ports, and seven of these involve two adjacent ports in Northern California: Crescent City and Fort Bragg. Only one vessel, which is in the group of seven, satisfies the minimum landings condition at three ports. Future work could include choice of port as another factor in the analysis.

The aggregate spatial distribution of effort associated with each port can be complicated, but is generally centered on a port, and therefore, transportation costs to a fishing location are likely to be roughly proportional to the straight line distance to that location, and fuel costs during tows roughly proportional to duration. Therefore, costs incurred traveling from port to the location of a tow are treated as a separate fixed factor at each port for areas inside and outside of the RCA. This assumption ignores variation in travel costs within each area, which tend to be smaller than traveling between areas due to the symmetry of effort around each port.

Fig. 2 shows ex-vessel prices, and total tow hours inside and outside of the RCA for DTS species in the twenty-one years between 1981-2001 at ten ports in California's groundfish trawl fishery. Disaggregated data from Fig. 2, for individual vessels, are used for analysis with the panel Tobit model. A separate analysis is conducted for each of the ten ports in Fig. 1. The plots in Fig. 2 show an interesting mix of relationships across ports. A downward trend in tow hours after about 1997 is noticeable across ports, which is caused by reductions in bimonthly landings limits for each vessel. Future work using monthly or bimonthly data could include these landings limits as a factor for individual vessels, and explicitly consider effects of seasonality.

Ex-vessel prices in Fig. 2 exhibit relatively little variation until about 1994, after which prices increase for a few years at all ports, and fluctuate thereafter. This rise in prices appears to be associated with an increase in tow hours at each port, both inside and outside of the RCA. There is no clear pattern of correlation between tow hours inside and outside of the RCA, but these do seem related at several ports. The panel Tobit model is suitable for estimating and testing dynamic relationships in fishing effort between areas, and for estimating how each series is influenced by ex-vessel prices.

4. Results

Results are presented in this section of applying the panel Tobit model in (2) to the trawl data described in the previous section. The first set of results compares likelihood values of the pooled model to the unrestricted model with individual heterogeneity using an asymptotic chi-squared likelihood ratio test (e.g. Amemiya, 1985). A separate test is conducted for each port in the data, under the assumption these are separate events. Any error or bias from this assumption is likely to be small, as described in the introduction.

The significance level for each chi-squared test is reported in Tab. 1. Degrees of freedom used to calculate the test statistics depends on numbers of vessels, which are given in the table. Ports in the table are ordered geographically from north to south. Restrictions imposed by the pooled model are rejected at the 5% level for Eureka, Fort Bragg, Bodega Bay, and San Francisco. For these ports, an additional likelihood ratio test was conducted using a cell corrected model that estimates α and β, a constant and coefficient on ex-vessel prices respectively, conditional on other parameters set equal to estimates from the pooled version of the model. This cell corrected model is not rejected at the ports listed above. A similar set of tests was conducted using a model with individual heterogeneity only in α, but this model has convergence problems, presumably because this parameter is not significant in many cases.

Tab. 2 reports maximum likelihood estimates for each parameter in the pooled model, and gives significance levels from a likelihood ratio test using the model with a zero value for each. Parameters are estimated for areas inside (R) and outside (O) of the RCA. Results in Tab. 2 show substantial variation across ports, but exhibit patterns consistent with economic theory. The coefficient on ex-vessel prices β is positive in all cases, and significant in most. Economic theory predicts an increase in prices should increase levels of fishing effort, and moreover, that prices should be a significant factor in decisions about fishing effort. Economic theory does not seem to have qualitative implications for the sign pattern of the coefficients γ or λ, but necessary conditions for stability of the model are these parameters have absolute values less than one. These conditions are satisfied in all cases with the pooled model. In addition, stability requires the product of λ values inside and outside of the RCA to be greater than the corresponding product with values of γ, which measure effects of past effort in one area on effort in the other. Point estimates for λ and γ do not satisfy this second condition for stability at two ports, Eureka and Bodega Bay, but in both cases at least one estimate of γ is not significant at the 5% level. In general, many of the coefficients on lagged effort are not significant at this level, which may suggest that dynamic relationships in fishing effort are not important, or at least less important than the dynamics of ex-vessel prices. However, further analysis is warranted before reaching this conclusion. Stability, or invertibility, conditions for the model also require the absolute value of the autocorrelation coefficient ρ to be less than one, which is satisfied by the pooled model in all cases.

A summary of maximum likelihood estimates and significance values are reported in Tab. 3 for the cell corrected model using data from the four ports where the pooled model was rejected. Estimates of α and β are computed for each individual conditional on values of other parameters set equal to point estimates from the pooled model. Presenting estimates for individual vessels is not practical, and may not be legal because the analysis in this paper uses confidential data. Therefore, only averages of the individual estimates are presented in Tab. 3, which may be compared with corresponding estimates from the pooled model in Tab. 2. Significance levels are computed for each parameter in Tab. 3 with a likelihood ratio test of zero values for all individuals. As with the pooled model, ex-vessel prices are significant in most cases. In two cases, the average coefficient on ex-vessel prices is negative, and is significant for San Francisco inside the RCA, which may imply the model is misspecified for some individuals at this port. Inspection of the estimates and significance tests for individual vessels at San Francisco shows the negative average value is caused by an estimate of β that is negative and significant for one vessel, and estimates for the other five vessels at this port are positive or not significant.

5. Policy Simulations

For results in this section, the fixed effects model in (1) and (2) includes a time series of regulatory data as an additional exogenous variable. The regression equations are modified to include a term with a coefficient that may express individual heterogeneity times an exogenous regulatory variable. As noted in the introduction, managers use bimonthly limits on landings to regulate trawlers in the West Coast groundfish trawl fishery. These limits apply to individual vessels and separate limits are established for a host of species. Data on trip limits are taken from the Pacific Fishery Management Council's Stock Assessment and Fishery Evaluation (SAFE) document for the 2001 fishery (PFMC, 2001). Separate trip limits exist for areas north and south of Cape Mendocino, California. Bi-monthly trip limits for each vessel were aggregated to form totals for each year, and these annual totals were used as the regulatory variable in the model. Separate limits on landings were considered for the target group of species used in this paper: dover soles, shortspine and longspine thornyheads, and sablefish; and several species of rockfish: darkblotched, bocaccio, canary, widow, and yellowtail. A series for lingcod was also constructed. Of these, species in the target group, bocaccio, and widow rockfish are most relevant.

Limits for bocaccio were implemented less than a decade ago, so there are few years of regulatory data, which is true for most of the species listed above. Trip limits for sablefish and widow rockfish go back to 1982, but limits for most other species do not start until after 1995, reflecting the time when problems for rockfish became apparent from stock assessments. Because trends in trip limits across rockfish species are similar, effects of these are confounded with those for widow rockfish. The same is true for dover sole and shortspine thornyheads, which are confounded with sablefish. Limits for longspine thornyheads do not become binding until 1998, and are not informative for this analysis. Therefore, limits for sablefish and widow rockfish were analyzed. Annual limits on landings of these species for each vessel in northern and southern areas are presented in Fig. 3. The northern limits apply to Crescent City and Eureka, and other ports are associated with the southern limits, which are not binding until 1995 for both species. The series for sablefish and widow rockfish look similar in some ways. However a policy to allow permit stacking, which is intended to proxy individual tradable quotas, is allowed for sablefish. Regressions using sablefish limits as a regulatory variable produce some conflicting results. Limits for widow rockfish produce reasonable results in most cases, which are presented below.

The first set of results from maximum likelihood simulations with regulatory variables are presented in Tab. 4, which gives significance levels from an asymptotic chi-squared likelihood ratio test based on a hierarchical nesting of individual heterogeneity among exogenous variables in three models: No heterogeneity (i.e. the pooled model), individual heterogeneity in the constant term and coefficient on ex-vessel prices, α and β in (2), and individual heterogeneity in α, β, and a coefficient on the regulatory variables, denoted by ζ. With policy variables included, the pooled model is rejected at the 5% level at six of the ten ports, compared to four ports in Tab. 1. This result could indicate that sensitivity to regulations is an important factor to consider for some individuals, but not others, which makes sense. Some vessels fish more than others. These vessels are more likely to have binding trip limits, and thus, are more sensitive to changes in regulations.

The pattern of rejections in Tab. 1 and Tab. 4 is mixed. For example, the pooled model is rejected for Fort Bragg in Tab. 1, but not in Tab. 4. In addition, the pooled model is rejected for Crescent City, Monterey, and Avila in Tab. 4, but it is not rejected for these ports in Tab. 1. These differences probably indicate the model without regulations suffers from an omitted variables bias that affects test results. Further differences with regulatory variables can be observed by comparing results for the model with individual heterogeneity in α and β in Tab. 1, which do not reject the model, to results for this case in Tab. 4, which reject the model for Crescent City, Eureka, and Fort Bragg. Moreover, even the policy model with individual heterogeneity in each of the exogenous variables, specifically in the coefficients α, β, and ζ, is rejected at two ports, which is different qualitatively from results in Tab. 1 that did not reject the model corrected to allow individual heterogeneity in the exogenous variables.

Maximum likelihood estimates and significance levels from the pooled model with policy variables are presented in Tab. 5, which are computed using the estimation and testing strategy described for results in the previous section. Like results in Tab. 4, parameter estimates and significance levels are sensitive in some cases to including regulatory variables. In two cases, significance at the 5% level of one or more variables changes relative to results in Tab. 2. The first case is Moss Landing, and values of γ and ρ are significant both inside and outside of the RCA in Tab. 2, but neither parameter is significant in Tab. 5. The other case is Avila, and again, values of γ inside and outside of the RCA are significant in Tab. 2 but not in Tab. 5. Note in the case of Moss Landing, the loss of significance in γ and ρ occurs even though coefficients on the regulatory variables, denoted by ζ, are not significant. Values of ζ, both inside and outside of the RCA, are not significant at another port, Bodega Bay, but regulations are a significant factor at the other eight ports.

Comparison of values in Tab. 5 to those in Tab. 2 shows numerous other cases where significance levels change with the addition of regulatory variables to the model. As above, results in Tab. 5 show the coefficient on ex-vessel prices β is positive and significant in most cases, but instead of the exception being Morro Bay, as it is in Tab. 2, ex-vessel prices are not significant at Avila in Tab. 5. Although gains and losses in significance for α and β are about the same, a value of γ inside or outside of the RCA is significant for five ports in Tab. 2, but only two ports in Tab. 5, implying that shifts in fishing effort between areas inside and outside of the RCA may be explained by effects of regulations, instead of dynamic interactions. Four more values of λ are significant in Tab. 5 than in Tab. 2, and three values of ρ that are significant in Tab. 2 are not in Tab. 5. In other words, including regulatory variables increases support for a model with simpler dynamics, although autocorrelation is significant at six ports in Tab. 5. Parameter values in Tab. 5 satisfy the stability conditions described above, with all point estimates of γ, λ, and ρ less than one in absolute value. Conditions on γ and λ that limit dynamic interactions between areas are trivially satisfied for all ports Tab. 5 because at least one value of γ is not significant at each.

Results in Tab. 6 give average maximum likelihood estimates and significance levels from the individual regressions at each port for which the pooled model is rejected, but the model corrected in coefficients α and β is not, according to Tab. 4. The point estimates of the coefficient on ex-vessel prices β are positive, and ex-vessel prices are a significant factor for Monterey and San Francisco, but not Avila, which is the same result for these ports as in Tab. 5. For Fort Bragg, the model corrected in coefficients α, β, and ζ is not rejected according to Tab. 4. Average maximum likelihood estimates and significance levels using the model that is corrected in these coefficients for Fort Bragg are also given in Tab. 6. In this case, results give less weight to ex-vessel prices on effort inside the RCA and the average value of point estimates for β in this area is negative, but not significant at the 5% level. However, negative point estimates for the coefficient on ex-vessel prices in this area occur for eight of the fourteen individuals at Fort Bragg, but the negative value is significant for only one in the individual regressions. Incidentally, ex-vessel prices are a significant effect inside the RCA for one other vessel, and the point estimate of β is positive in this case.

Result for Fort Bragg in Tab. 6 may be more difficult to explain, which is the negative and significant values of ζ. Taken literally, this result implies that relaxing trip limits on widow rockfish decreases expected fishing effort, which is at odds with results for other ports in Tab. 5 that indicate in each case where ζ is significant, it is also positive. Results for Fort Bragg may be explained by a decrease in trip limits for widow rockfish shifting trawl effort from targeting rockfish to DTS species. In other words, the experience at other ports could be that decreases in limits on widow rockfish constrain effort for DTS species because of bycatch considerations, but at Fort Bragg, widow rockfish are an alternative target strategy, and a decrease in limits for widow rockfish could shift effort to DTS species. Results using limits on sablefish, instead of limits on widow rockfish, as the regulatory variable in the model show that average point estimates for ζ are positive, but not significant. In addition, estimates of α and β are not significant in the model for Fort Bragg with limits on sablefish, further demonstrating the sensitivity of the model to the regulatory variables.

6. Discussion

This paper uses a panel Tobit model, based on the panel models in Lee (1998) and the dynamic Tobit models in Lee (1999), to compute maximum likelihood estimates and test statistics for individual groundfish trawlers that have operated at ten ports in California. The model includes autocorrelation in the disturbances, which complicate the likelihood function with high-dimensional integrals. The analysis reported in this paper follows the simulated maximum likelihood approach developed by Lee, which is based on a version of the Gibbs sampler that accomodates censored data.

In addition to parameter estimation, an analysis of covariance was conducted at each port to test whether parameter restrictions across equations for individual vessels are significant. These restrictions are necessary to obtain unbiased estimates when using pooled or aggregated data for analysis, which is done frequently in applied time series analysis, including fisheries (e.g. Dalton, 2001; Dalton and Ralston, 2004). The principal objective of this paper is to provide a framework for testing hypotheses about the use of pooled data to determine whether results of past work may be biased, and if so, to guide future model development. Likelihood ratio tests were conducted by comparing likelihood values of a pooled model, with each parameter restricted to be equal across vessels, to an unrestricted dynamic model with a separate set of parameters for each vessel. In cases where parameter restrictions necessary for unbiased estimates with pooled data are rejected, a cell corrected version of the model was tested that allows coefficients on ex-vessel prices, an exogenous variable, to vary across vessels. A version of the model with a second exogenous variable, limits on landings of widow rockfish, was also tested.

Results in the paper imply the relatively simple spatial and autoregressive model of fishing effort captures the main features of the groundfish trawl data, including effects of ex-vessel prices and past effort on current decisions about where to fish, and how much effort to deploy. Maximum likelihood estimates are consistent with predictions of economic theory, for example showing positive responses in fishing effort to price increases. In addition, the estimated coefficients on lagged variables are consistent with stability conditions for the model. Results show that ex-vessel prices are usually a significant variable, which economic theory implies, but the pattern of significance for other parameters is mixed. Past effort is significant at the 5% level for most ports in the analysis, implying that dynamic variables are important to consider in decisions about fishing effort. Autocorrelation in the model's error process is also significant at most ports, implying that a straightforward maximum likelihood approach with the panel Tobit model would not be appropriate. For the model with ex-vessel prices as the only exogenous variable, test results show that parameter restrictions associated with using pooled data are not rejected for six ports, and the cell corrected model that allows individual variation in responses to ex-vessel prices is not rejected at the other four. This result implies that using simplified models, based on data that are at least partially aggregated, to analyze behavior in the groundfish trawl fishery would be acceptable practice. However, this result is sensitive to including regulations as an exogenous variable.

The policy version of the model includes limits on landings of widow rockfish as an exogenous variable. Limits on widow rockfish are significant at the 5% level at eight of the ten ports that were analysed. At seven of these, a reduction in the limit on widow rockfish implies a decrease in fishing effort for DTS species, the target group used for analysis in the paper, which is reasonable because rockfish bycatch is associated with trawling for DTS species. Results for the eighth port, Fort Bragg, could make sense if targeting widow rockfish is an alternative to trawling for DTS species, which was true at least prior to 1990. Using trip limits for sablefish, one of the DTS species, instead of widow rockfish, as the regulatory variable for Fort Bragg yields estimates that imply a reduction in trip limits reduces effort, although the estimates are not significant. More generally, test results with the policy model show that parameter restrictions associated with using pooled data are not rejected for only four ports. The model corrected for individual responses to ex-vessel prices is not rejected for three ports, and the model corrected for individual responses to ex-vessel prices and regulations is not rejected for one. Therefore, an analysis of covariance using the policy version of the model cannot rule out significant individual effects in the panel Tobit model's dynamic variables at two ports.

On the other hand, including regulatory variables in the model clarifies or sharpens dynamic relationships to some extent. For example, the number of ports where autocorrelation is significant falls from eight to six in the policy version of the model. The number of ports where lagged effort is significant rises from six to nine, and significant dynamic interactions between areas inside and outside of the Rockfish Conservation Area (RCA) falls from four ports to two. An explanation for these changes is provided by the spatial pattern of significance for the regulatory variables. In most cases, limits for widow rockfish are a significant factor inside the RCA, which makes sense because these were designed specifically to constrain effort to areas with low bycatch of rockfish species. In other words, reducing trip limits for widow rockfish tends to encourage DTS trawlers to move to areas with less bycatch, which is the area outside of the RCA. In the model without regulatory variables, these shifts in effort are attributed to dynamic factors, but in fact, are the effects of regulations. These results also point to a limitation of trip limits, which may discourage fishing in an area, but in general, these limits will not eliminate fishing effort there. Results for Fort Bragg indicate that reducing trip limits for rockfish may have even increased trawl effort inside the RCA for DTS species. So, the RCA is probably needed to manage bycatch.

Currently, fishery managers are lacking reliable statistical tools to analyze or predict dynamic changes in fishing effort in response to spatial closures such as the RCA. The dynamic and spatial model in this paper is an appropriate framework for analyzing these closures. The model of fishing effort used in this paper may be interpreted as an unrestricted version of a bioeconomic model under an information structure based on rational expectations. In future work, the cross-equation parameter restrictions implied by rational expectations may be used to test this information structure against alternatives, which has been done in previous work using pooled time series data. The advantage of an approach based on rational expectations is to identify structural parameters that, at least in theory, are invariant to changes in fisheries management, including spatial closures such as the RCA.

Results in this paper imply the structural fisheries models that have been used in past work should be extended to allow individual responses to ex-vessel prices, regulations, and even dynamic variables. These modifications will enhance the equilibrium structure of the models used in previous work, which have been based on a symmetric rational expectations competitive equilibrium in which all individuals are assumed to be identical. Moreover, extending the structural models to allow for individual responses to ex-vessel prices will guide further development of the econometric model in this paper, as an unrestricted alternative to the structural models. For example, the panel Tobit model in this paper could be extended to include, at least, a dynamic equation for ex-vessel prices, so the model would be a vector autoregression for panel data (Holtz-Eakin, Newey, and Rosen, 1989; Hsiao and Pesaran, 2001), but with censored endogenous variables. Other possibilities for future work are noted in the paper, such as more refined spatial and temporal scales, and comparing model predictions from data through 2001 with more recent data.

Results in this paper suggest the future use of panel Tobit models in fisheries management is promising. Future empirical work could include the application of panel Tobit models to groundfish trawlers in Oregon and Washington, where panel data like those used for trawlers in California are available. In addition, use of panel Tobit models to analyze halibut and large-scale groundfish, and crab, fisheries in the Gulf of Alaska and Eastern Bering Sea are possible. Structural models for these fisheries would provide econometrically estimated decision rules to analyze implications of spatial management, and future applications could include fishery rationalization programs that involve market-based mechanisms such as individual quotas.

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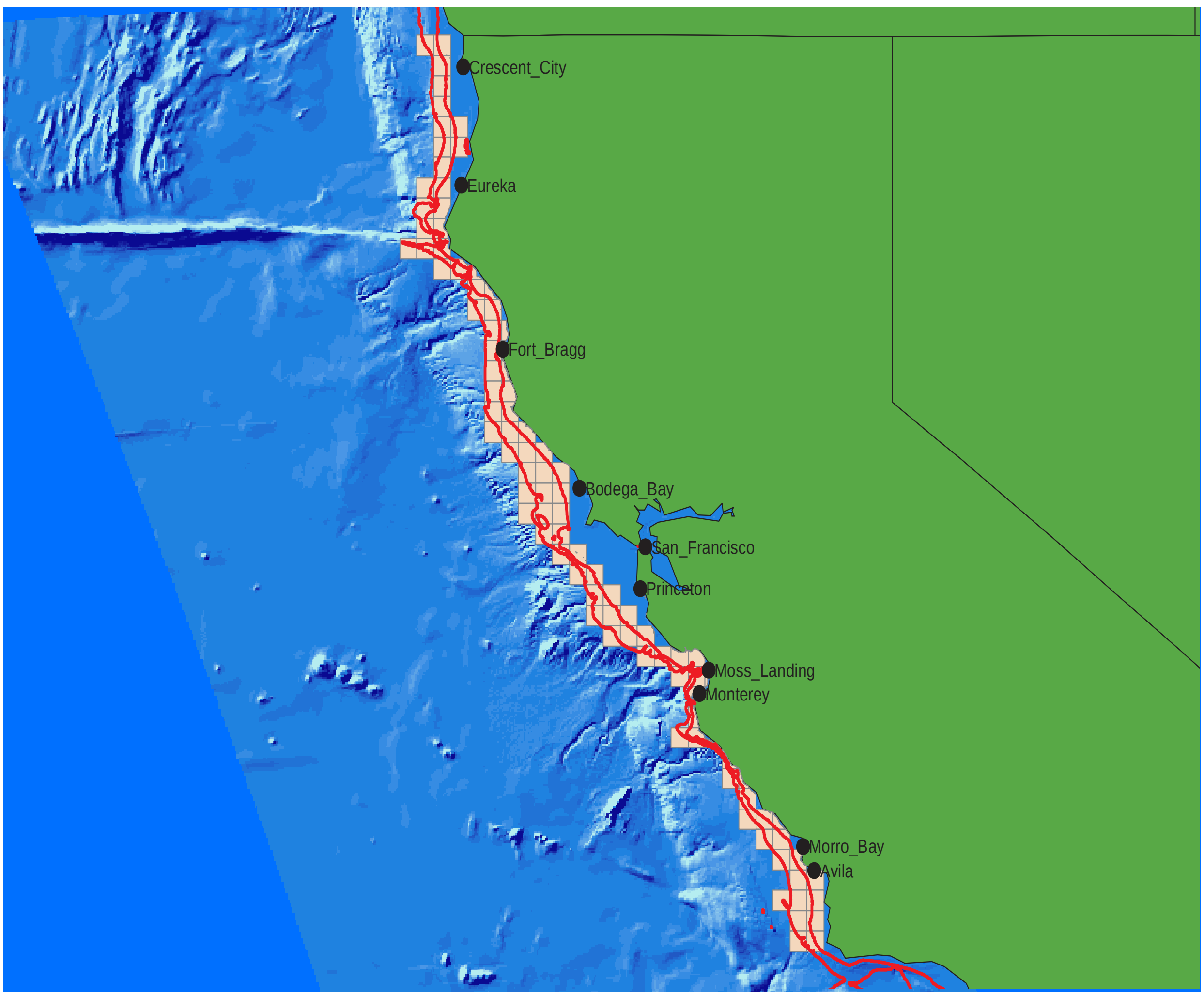

Figure 1: California Rockfish Conservation Area and Trawl Ports.

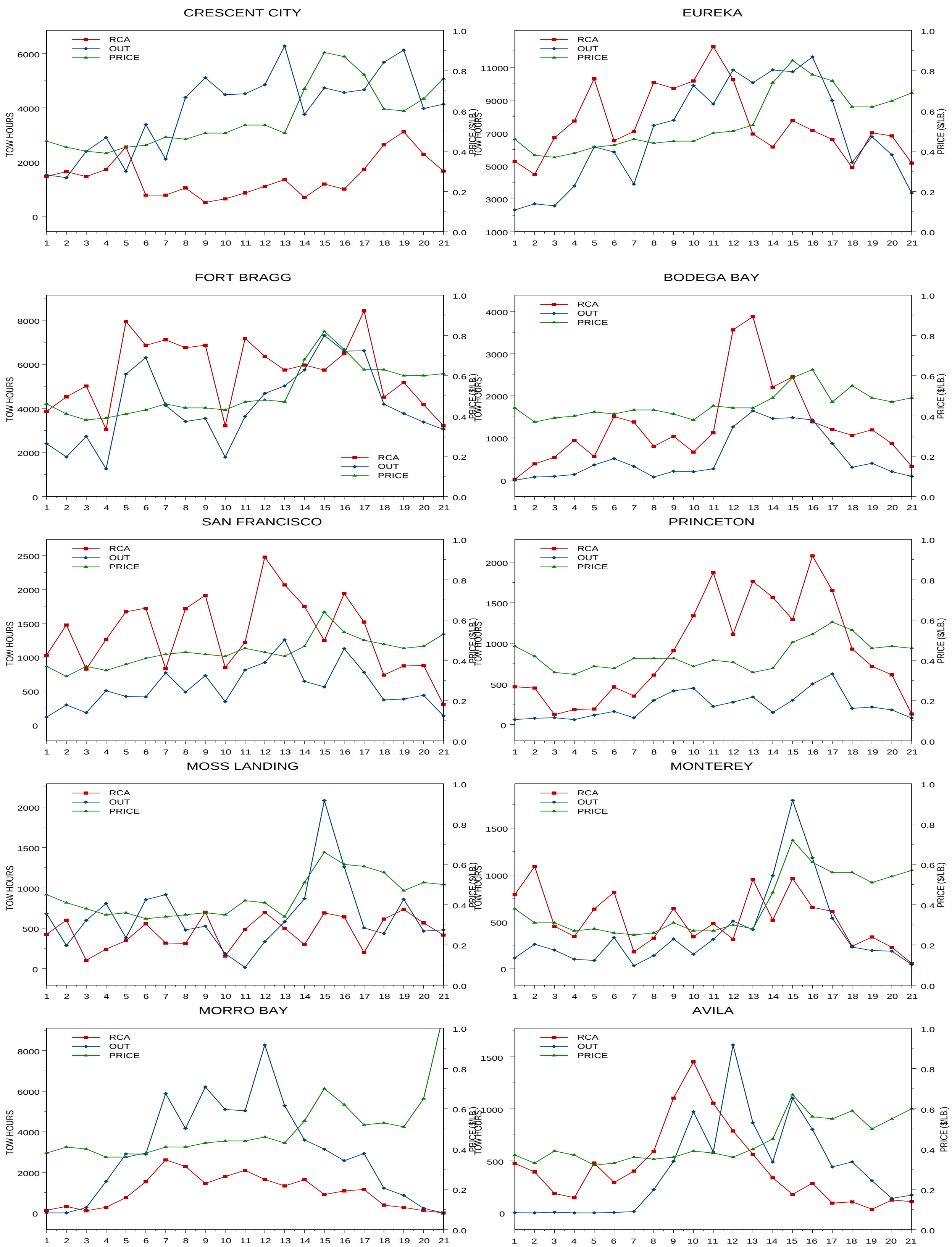

Figure 2: Port Total Tow Hours and Average Ex-Vessel Prices, 1981-2001.

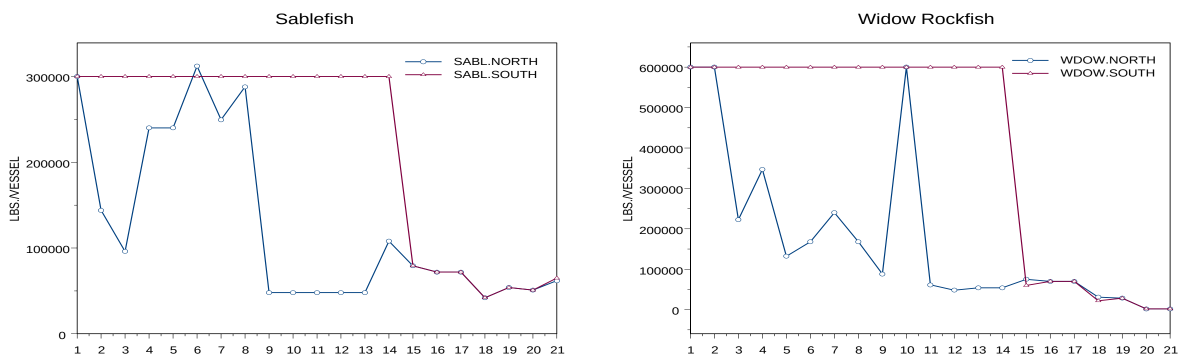

Figure 3: Annual Per Vessel Limits on Sablefish and Widow Rockfish Landings, 1981-2001.

Table 1: Significance Levels from Asymptotic Chi-Squared Likelihood Ratio Tests of Pooled and Cell Corrected Models.

Port	N	Pooled Model	Cell Corrected Model
Crescent City	22	0.982	
Eureka	37	0.040	0.770
Fort Bragg	14	0.000	0.489
Bodega Bay	7	0.000	0.955
San Francisco	6	0.002	0.302
Princeton	8	0.198	
Moss Landing	6	0.582	
Monterey	3	0.119	
Morro Bay	9	1.000	
Avila	4	0.793	

Table 2: Maximum Likelihood Estimates and Asymptotic Chi-Squared Significance Levels for the Pooled Model in Areas Inside (R) and Outside (O) of the RCA.

Port	RCA	α	sig.	β	sig.	γ	sig.	λ	sig.	ρ	sig.	σ
Crescent City	O	-22.11	0.000	232.74	0.000	-0.113	0.066	0.613	1.000	-0.076	0.308	190.40
	R	-6.46	1.000	85.06	0.000	-0.009	1.000	0.536	0.375	0.031	1.000	75.14
Eureka	O	6.35	1.000	454.43	0.000	0.123	1.000	-0.184	1.000	0.748	0.550	183.34
	R	13.31	0.289	400.10	0.000	-0.036	0.243	0.084	0.025	0.641	0.000	174.90
Fort Bragg	O	12.55	0.129	176.14	0.007	-0.041	1.000	0.771	0.362	-0.025	0.339	213.71
	R	81.31	0.002	115.81	0.099	0.035	0.183	0.695	0.000	-0.267	0.001	245.99
Bodega Bay	O	1.91	0.916	107.14	0.003	0.107	1.000	0.020	0.053	0.535	0.016	72.46
	R	0.74	0.487	276.79	0.000	0.397	0.041	0.274	0.053	0.353	0.000	146.34
San Francisco	O	2.84	1.000	42.55	1.000	0.145	0.033	0.490	0.004	0.003	1.000	95.77
	R	-14.43	1.000	513.94	0.001	0.065	0.385	0.172	1.000	0.514	0.016	208.99
Princeton	O	-2.54	0.848	54.24	0.008	-0.008	0.165	0.419	0.771	0.001	0.630	46.59
	R	19.17	0.411	56.93	0.227	0.088	0.726	0.703	0.034	-0.222	0.036	124.43
Moss Landing	O	-59.96	0.001	391.10	0.000	-0.256	0.000	0.565	0.000	-0.098	0.000	126.09
	R	-30.13	0.000	194.23	0.000	0.016	0.000	0.468	0.000	-0.175	0.000	80.54
Monterey	O	-17.91	0.568	414.43	0.002	0.066	0.764	0.055	0.530	0.542	0.055	120.69
	R	-6.45	0.132	143.95	0.108	-0.264	0.061	0.905	0.021	-0.437	0.011	134.76
Morro Bay	O	49.73	0.001	142.41	0.355	0.213	1.000	0.609	0.220	0.057	0.184	269.82
	R	20.26	0.001	22.49	1.000	0.003	1.000	0.757	1.000	-0.115	0.304	115.59
Avila	O	-43.24	0.704	200.62	0.149	0.126	0.020	0.605	0.734	-0.393	0.012	158.02
	R	-30.42	0.603	190.81	0.035	-0.164	0.010	0.824	0.117	-0.227	0.272	114.02

Table 3: Average Maximum Likelihood Estimates and Asymptotic Chi-Squared Significance Levels for the Cell Corrected Model in Areas Inside (R) and Outside (O) of the RCA.

Port	RCA	N	α	sig.	β	sig.
Eureka	O	36	12.89	0.199	400.97	0.000
	R	36	16.62	0.150	429.84	0.000
Fort Bragg	O	14	79.85	0.063	62.50	0.012
	R	14	228.26	0.007	-88.95	0.103
Bodega Bay	O	7	0.16	0.973	116.22	0.112
	R	7	4.94	0.991	302.06	0.000
San Francisco	O	6	18.63	0.883	11.48	0.843
	R	6	103.25	0.004	-7.70	0.001

Table 4: Significance Levels from Asymptotic Chi-Squared Likelihood Ratio Tests of Models with Individual Responses to Exogenous Variables.

Port	N	Pooled Model	Corrected in (α, β)	Corrected in (α, β, ζ)
Crescent City	22	0.002	0.006	0.012
Eureka	37	0.000	0.001	0.001
Fort Bragg	14	0.000	0.000	0.070
Bodega Bay	7	0.083		
San Francisco	6	0.000	0.071	
Princeton	8	0.097		
Moss Landing	6	0.130		
Monterey	3	0.029	0.090	
Morro Bay	9	1.000		
Avila	4	0.009	0.116	

Table 5: Maximum Likelihood Estimates and Asymptotic Chi-Squared Significance Levels for the Pooled Policy Model in Areas Inside (R) and Outside (O) of the RCA.

Port	RCA	α	sig.	β	sig.	ζ	sig.	γ	sig.	λ	sig.	ρ	sig.	σ
Crescent City	O	-61.1	0.000	248.1	0.000	44.2	0.027	-0.096	1.000	0.621	0.000	-0.094	1.000	188.14
	R	-7.8	1.000	80.2	0.000	0.7	1.000	-0.015	1.000	0.613	0.054	-0.011	1.000	74.91
Eurkea	O	-162.6	1.000	368.8	0.000	120.6	1.000	-0.076	1.000	0.658	1.000	-0.108	1.000	183.90
	R	-94.0	0.031	443.7	0.000	132.4	0.000	-0.054	0.173	0.400	0.001	0.313	0.000	173.45
Fort Bragg	O	53.2	1.000	203.6	0.012	-209.5	0.004	-0.014	1.000	0.652	1.000	0.192	0.147	211.23
	R	92.4	0.000	52.5	0.000	-367.8	0.000	0.099	0.000	-0.257	0.000	0.811	0.000	228.29
Bodega Bay	O	7.7	1.000	108.5	0.001	-102.4	1.000	0.064	0.164	0.222	1.000	0.395	0.033	72.43
	R	2.6	1.000	287.2	0.000	22.1	1.000	0.569	0.063	0.134	0.027	0.371	0.000	147.46
San Francisco	O	-40.7	1.000	83.6	0.087	41.0	0.408	0.136	0.003	0.506	0.065	-0.029	1.000	94.40
	R	-107.5	1.000	781.0	0.000	315.1	0.006	0.111	0.275	-0.021	1.000	0.637	0.321	200.76
Princeton	O	-10.4	0.289	59.6	0.004	9.6	0.220	-0.006	0.729	0.416	0.079	-0.006	0.772	46.46
	R	-43.9	0.116	101.6	0.036	50.7	0.010	0.026	0.801	0.727	0.002	-0.266	0.008	121.94
Moss Landing	O	-86.4	0.005	442.7	0.000	12.0	0.499	-0.258	0.214	0.541	0.003	-0.091	0.312	126.39
	R	-28.3	0.381	194.8	0.000	-3.9	0.154	0.006	1.000	0.482	0.020	-0.184	0.470	80.17
Monterey	O	-15.2	0.117	431.1	0.001	-16.5	0.743	0.092	0.736	-0.004	0.551	0.576	1.000	120.26
	R	-119.2	0.009	209.8	0.007	80.2	0.012	-0.208	0.261	0.877	0.001	-0.488	0.002	128.29
Morro Bay	O	-219.7	0.046	351.8	0.008	258.6	0.000	0.154	1.000	0.607	0.054	-0.055	0.177	255.59
	R	-104.6	0.197	125.4	0.120	101.0	0.000	-0.029	1.000	0.789	0.048	-0.172	0.636	110.89
Avila	O	-17.0	1.000	267.1	0.172	-54.5	0.073	0.203	1.000	0.479	0.012	-0.261	0.007	162.52
	R	-44.5	1.000	131.4	0.065	38.6	0.046	-0.115	0.053	0.816	1.000	-0.219	0.272	113.77

Table 6: Average Maximum Likelihood Estimates and Asymptotic Chi-Squared Significance Levels for Policy Model Corrected in (α, β) or (α, β, ζ) for Areas Inside (R) and Outside (O) of the RCA.

Port	RCA	N	α	sig.	β	sig.	ζ	sig.
Monterey	O	3	-72.4	0.256	549.8	0.000		
	R	3	-181.4	0.000	281.9	0.000		
San Francisco	O	6	-3.2	0.879	36.5	0.772		
	R	6	-48.4	0.000	447.8	0.000		
Avila	O	4	-29.6	0.912	28.9	0.269		
	R	4	-47.0	0.354	88.9	0.289		
Fort Bragg	0	14	130.5	0.001	57.0	0.004	-256.5	0.005
	R	14	142.4	0.000	-333.4	0.072	-359.6	0.000