

What are we protecting? Fisher behavior and the unintended consequences of spatial closures as a fishery management tool

JOSHUA K. ABBOTT^{1,3} AND ALAN C. HAYNIE²

¹*School of Sustainability, Global Institute of Sustainability and ecoSERVICES Group, Arizona State University, P.O. Box 875502, Tempe, Arizona 85287 USA*

²*Resource Ecology and Fisheries Management Division, Alaska Fisheries Science Center, National Marine Fisheries Service, NOAA, 7600 Sand Point Way NE, Building 4, Seattle, Washington 98115 USA*

Abstract. Spatial closures like marine protected areas (MPAs) are prominent tools for ecosystem-based management in fisheries. However, the adaptive behavior of fishermen, the apex predator in the ecosystem, to MPAs may upset the balance of fishing impacts across species. While ecosystem-based management (EBM) emphasizes the protection of all species in the environment, the weakest stock often dominates management attention. We use data before and after the implementation of large spatial closures in a North Pacific trawl fishery to show how closures designed for red king crab protection spurred dramatic increases in Pacific halibut bycatch due to both *direct* displacement effects and *indirect* effects from adaptations in fishermen's targeting behavior. We identify aspects of the ecological and economic context of the fishery that contributed to these surprising behaviors, noting that many multispecies fisheries are likely to share these features. Our results highlight the need either to anticipate the behavioral adaptations of fishermen across multiple species in reserve design, a form of implementation error, or to design management systems that are robust to these adaptations. Failure to do so may yield patterns of fishing effort and mortality that undermine the broader objectives of multispecies management and potentially alter ecosystems in profound ways.

Key words: bycatch; ecosystem-based management; fisher behavior; implementation error; marine protected areas; multispecies management; spatial closures.

INTRODUCTION

Spatial fishery closures such as marine protected areas (MPAs) are an increasingly prominent tool for marine resource management. The potential benefits of these policies include habitat protection (Turner et al. 1999), increased biodiversity and biomass within reserve boundaries (Lester et al. 2009), the protection of vulnerable species and life stages (Hooker and Gerber 2004), reductions in bycatch (Hobday and Hartmann 2006), buffering against uncertainty (Allison et al. 2003, Grafton et al. 2005), and potential yield spillovers to exploitative uses in open areas (Roberts et al. 2001, Gell and Roberts 2003). MPAs are often upheld as a key tool for successful ecosystem-based management (EBM) of fisheries (Witherell et al. 2000, Pikitch et al. 2004, Roberts et al. 2005) and an essential element of sustainable fisheries policy (Pauly et al. 2002, Worm et al. 2009). They have received considerable support from advocacy groups and are often defined as a worthy goal in themselves for the sustainable development of coastal nations (United Nations 2002).

Nevertheless, spatial closures remain controversial as a fishery management tool. Their success is predicated

upon ample monitoring and enforcement, and economic, institutional, and biological factors determine whether their net benefits are positive to fishermen and other stakeholders (Sanchirico and Wilen 2001, Sanchirico 2005). There is typically also uncertainty over the magnitude, timing, and distribution of costs and benefits (Smith et al. 2010). One aspect of closures that has received some attention is the role of adaptive fisher behavior in response to spatial management measures, a subset of a larger literature focusing on "fleet dynamics" more generally (e.g., Salas and Gaertner 2004, Branch et al. 2006; Putten et al., *in press*). In addressing this question empirically, economists have primarily focused on developing empirical models that can be used to understand and predict the patterns of substitution over space that result from the implementation of spatial constraints (Curtis and Hicks 2000, Smith and Wilen 2003, Haynie and Layton 2010, Zhang and Smith 2011). Fisheries scientists have also recognized the important role of fisher behavior in conditioning the biological outcomes of marine reserves (Botsford et al. 2003, Halpern and Warner 2003, Halpern et al. 2004). Studies of effort dynamics with respect to spatial closures have often been examined in the context of measuring medium to long-term fishery spillovers from marine reserves (McClanahan and Mangi 2000, 2001, Goni et al. 2008) with significant effort specifically addressing the phenomenon of "fishing the line," the areas

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³ E-mail: Joshua.K.Abbott@asu.edu

immediately outside closures (Wilcox and Pomeroy 2003, Kellner et al. 2007). A common finding of the biological and economics literatures is that failure to adequately consider fisher behavior when designing management policies is a form of implementation error (Rosenberg and Restrepo 1994) and can severely undermine the success of fisheries management measures (Wilen et al. 2002).

The complexities associated with fishers' behavioral responses to spatial management measures are further compounded when fishers target, or have the capacity to target, a wide array of species and managers are tasked with balancing impacts across species (as in EBM). Fishermen are often extremely effective at adapting to closures in an attempt to maximize their profits. The displacement of fishing to areas that remain open can create new management challenges in that, while many species may share affinities of habitat that make their joint protection within a closure effective, these spatial complementarities may not always arise. If species exhibit distinct patterns of spatial abundance, then efforts to protect one species through spatial measures can inadvertently increase harvesting pressure on another. Depending on context, this displacement may be more or less desirable under EBM, but certainly requires careful consideration in management planning. This planning may be made significantly more difficult by the potential for large-scale spatial management measures to trigger complex behavioral changes by fishermen that go beyond the mere spatial reallocation of effort. In particular, closures may trigger changes in the targeting behavior of fishermen with potential cascading effects for bycatch species as well. In this complex multispecies context, it is unclear whether the expected ecosystem benefits from MPAs and other forms of spatial management will necessarily outweigh the short- and long-run costs that stem from the behavioral adaptation of fishermen.

To illustrate the real-world potential of these concerns, we examine two major permanent spatial closures in the U.S. Eastern Bering Sea (EBS) using high-resolution vessel-level observer data from before and after their implementation. Vessels in this bottom-trawl fishery target rock sole (*Lepidopsetta polyxystra*) roe and Pacific cod (*Gadus macrocephalus*) but also experience bycatch of red king crab (*Paralithodes camtschaticus*; RKC) and Pacific halibut (*Hippoglossus stenolepis*), which are "prohibited species catch" (PSC) that must be discarded. Bycatch has historically served as a binding constraint on fishing for the target species and fishing has been seasonally curtailed when the fleetwide allocation of either PSC species is reached. RKC bycatch has been of particular concern given the depressed state of its stocks relative to historical highs and the possible role of trawl bycatch in this decline (Dew and McConnaughey 2005), while halibut is highly valued as a target species that benefits many Alaska

coastal communities. In 1995, a large (~13 700 km²) MPA, known as the Red King Crab Savings Area (RKCSA), that prohibits all bottom trawling within its boundaries and was designed to protect aggregations of female RKC and their habitat was established (Fig. 1). Analysis by National Marine Fisheries Service (NMFS) found that 40–70% of RKC bycatch in the fishery came from the RKCSA in the years prior to the closure (NMFS 1995). A second no-trawl reserve, the Pribilof Islands Habitat Conservation Area (PHCA), was simultaneously implemented as a sanctuary for broad conservation purposes but with special preference for protecting depressed stocks of non-RKC crab, marine mammals, and seabirds.

Our analysis makes several contributions. First, we extend the literature on the pre/post-reserve comparisons of fishing effort and catch (e.g., Murawski et al. 2005) by utilizing a data set of unprecedented consistency and detail from before and after a closure event. Second, we are able to isolate indirect behavioral effects of the closure on catch composition through induced shifts in the targeting behavior of fishermen. Such behavior may have consequences for bycatch as well. We show how the combination of direct spatial displacement and indirect behavioral effects led to substantial additional harvest pressure on Pacific halibut. Third, we examine the ecological, technical, and economic factors that made the closures vulnerable to behavioral adaptations and argue that these are features shared by many multispecies fisheries.

Our focus is on identifying and understanding the behavior of fishermen within the management system, not attempting to uncover the long-term performance of the closures in our particular case. While very important, the complex dynamics of the ecosystem and the large variation in recruitment of many populations make the latter a formidable task, and any assessment would require grappling with considerable uncertainty in the assessments of less studied species. Instead, our objective is to glean lessons about the behavior of fishermen as the apex predator in the system, lessons that may prove essential in the effective management of many other multispecies fisheries.

Our investigation is organized around four key questions. First, how did the distribution of fishing effort change in the wake of the implementation of the closures? Second, did the closures (particularly the RKCSA) achieve the primary, single-species fishery management objective of reducing red king crab bycatch? Third, did the advent of closures measurably alter the targeting behavior of fishermen apart from the direct effect of spatially shifting their effort? Finally, how did the bycatch of halibut, another prohibited species, change as a result of the direct effect of displacement of fishing effort from the closures vs. indirect effects due to closure-induced changes in targeting behavior?

METHODS

Data

The primary data for this analysis are observations on the location and catch of each vessel from the North Pacific Groundfish Observer Program (NPGOP). While these data are confidential, a description of the observer program and a summary of the data it collects are publically available (database *available online*).⁴ During our study period, vessels 125 feet (38.1 m) and longer were required to carry an observer on all fishing days, while smaller vessels were only required to carry an observer on 30% of days and had discretion over when to satisfy these coverage requirements. Since onboard observers gather general data for each haul of the net (including date, tow duration, and total catch), we observe essentially the entire fishing history of vessels with 100% observer coverage. Observers randomly select a proportion of observed hauls for species composition sampling, including prohibited species catch (PSC) species like RKC and halibut. In these cases, a portion of the total catch is randomly selected in order to provide statistically robust estimates of the species composition of the haul.

The data collected by NPGOP observers is widely held to be of high quality, and is extensively used in research, stock assessment, and day-to-day management contexts. Observer debriefings after the termination of cruises and electronic logic checks help ensure the integrity of the data, and there are substantial criminal penalties for deliberate data fouling on the part of crew or observers. While observers are not enforcement agents for NMFS, their presence and the information they provide is critical to enforcement, ensuring that policies such as the counting and mandatory discard of PSC species are strictly enforced. An independent review of the NPGOP found that it compared favorably in both design and implementation to other observer programs (MRAG Americas 2000).

We focus on active participants in the rock sole fishery from January to March of 1992–1997; this provides us with three years of data on either side of the establishment of the closures. We limit ourselves to this relatively narrow window of time in order to focus on the *direct* behavioral responses of fishermen (and the associated impacts on catch and bycatch) to the closures while limiting concerns that our assessments of these direct effects are contaminated by long-run feedbacks through stock dynamics. We constrain the analysis to the early-season rock sole and cod fishery since this fishery exhibited significant bycatch of both red king crab and halibut and was the primary target of the RKCSA closure. To focus our sample on only those vessels participating in this fishery, we exclude all hauls in the Aleutian Islands management area (where a distinct fishery for Atka mackerel is pursued) and use

weekly production data to exclude all hauls in a week by vessels that failed to devote at least 50% of their weekly production to rock sole or Pacific cod or that had zero rock sole production. The latter criterion is designed to eliminate vessels that retained significant cod from the distinct Aleutian fishery (where rock sole is not a target). We further limit our sample to exclude hauls on days that the retention of cod or rock sole is prohibited or curtailed due to PSC-related management measures. Finally, we eliminate six vessels with vessel lengths less than 125 feet due to concerns that partially observed data from this portion of the fleet will not be representative since owners of partially observed vessels control the timing of observer coverage. Altogether, we have approximately 10 300 observations with 44% of these containing data on the species composition of catch.

In addition to observer data, we also utilize annual data on the biomass of rock sole, cod, red king crab, and halibut. Estimates of historical biomass on the Eastern Bering Sea shelf for rock sole and cod are taken from the 2005 stock assessment reports prepared by NMFS for the North Pacific Fisheries Management Council (Thompson and Dorn 2005, Wilderbuer and Nichol 2005) and are based upon fishery-independent annual groundfish bottom trawl surveys. Estimates of red king crab biomass are likewise derived from NMFS trawl surveys and record estimates of crab biomass in Bristol Bay and the Pribilof Islands region, spanning the rock sole fishing grounds (Rugalo et al. 2006). While stock assessment for Pacific halibut is the task of the International Pacific Halibut Commission (IPHC), coverage of the Eastern Bering Sea by setline surveys was relatively sparse at the time of this study and assessment relies heavily upon data gathered during NMFS groundfish trawl surveys. For this reason, and because trawl surveys are more likely than setline surveys to parallel the selectivity of the trawl fishing fleet, we use swept-area estimates of halibut abundance from the EBS groundfish trawl survey, summing estimates across all EBS strata to produce an overall index of annual biomass (Lauth 2011).

We also use data on the production and product value of catcher-processor vessels. Production quantity data come from mandatory weekly production reports detailing the weekly output of each species/product combination and the estimated pre-processed mass of the catch utilized for each product. We combine these data with estimates of fishing duration from the observer data to calculate the raw (unprocessed) rock sole and cod catch that is retained for production per hour of weekly fishing. Estimates on value of production are derived from annual reports filed by catcher processors. These data are used to calculate production-weighted annual prices for the dominant final products for rock sole (roe-in headed and gutted product) and cod (“eastern cut” headed and gutted). Both the weekly

⁴ http://www.afsc.noaa.gov/FMA/fma_database.htm

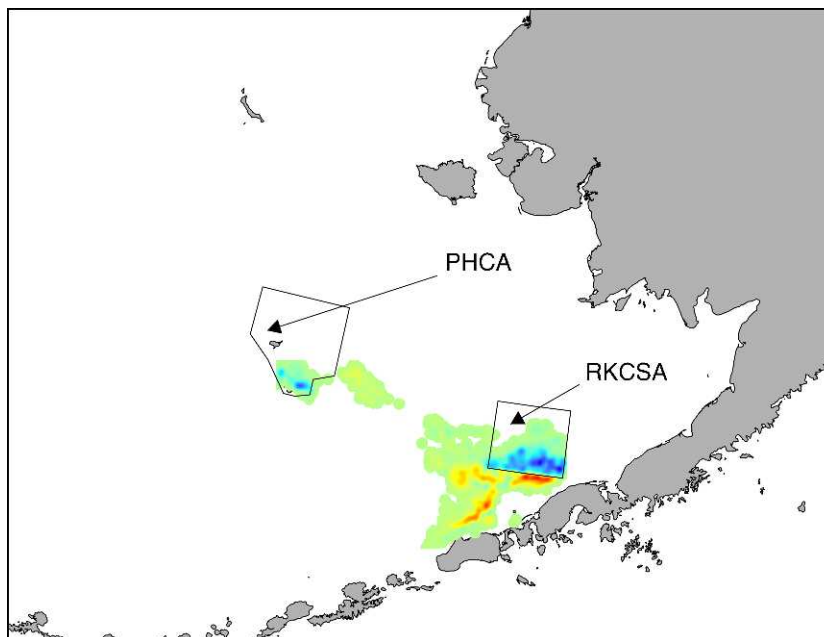


FIG. 1. The difference between 1995–1997 and 1992–1994 kernel-smoothed spatial densities of observed trawls in the January–March rock sole/cod fishery. Dark blue areas indicate regions with substantial relative loss of effort from 1995 onward relative to previous years, while orange areas indicate substantial gains. Green areas indicate no significant change in density, and white areas indicate regions with no fishing in either period.

production reports and annual reports are confidential economic data.

Models

Question 1: How did fishing effort redistribute in the wake of the closures?—To provide a visual assessment of the shift in fishing effort before and after the closures, we construct two distinct kernel density plots of the spatial density of hauls from 1992–1994 and 1995–1997. Each plot is defined as a probability density so that they are directly comparable in spite of different numbers of total observations in the two time intervals. Each plot was constructed using the kernel density command in ESRI ArcToolbox in ArcMap 10.0 (ESRI, Redlands, California, USA). A quadratic kernel was employed (Silverman 1998) using a relatively small search radius of 10 km in order to retain significant detail in the spatial intensity of hauls. Experimentation with larger and smaller bandwidths showed little discernible difference in the visual results. The 1992–1994 density is then subtracted from the 1995–1997 density to yield a summary of the change in the probability of fishing at any point.

To measure the catch and bycatch rate for each haul, we calculate the raw CPUE (kg/h) for each target and PSC species (rock sole, Pacific cod, Pacific halibut, and red king crab, respectively) by dividing the observer estimate of catch by the duration of the haul. To investigate how shifts in the distribution of fishing effort conform with spatial patterns of target and PSC catch intensity, we draw upon local measures of spatial

association (Getis and Ord 1992, Anselin 1995). Preliminary analysis using local Moran's I statistics revealed that patterns of negative spatial correlation in catch rates were not present in our data for any of the catch or bycatch species. We therefore focus on understanding the patterns of clustered catch rates in our data. To summarize these patterns, we use the G_i^* statistic (Getis and Ord 1992). Unlike the local Moran's I , the G_i^* statistic presupposes nonnegative spatial autocorrelation (i.e., clustering rather than repulsion of catch rates), but it has the advantage of being able to differentiate patterns of clustering of high and low catch rates. The statistic compares the proximity-weighted sum of catch rates at a particular point to the sum across the entire sample to construct a Z statistic where large values indicate “hot spots” of unusually high values, and low values indicate “cold spots.” To construct the sum around each point we used a “zone of indifference” weighting that ascribes equal weight to all hauls within a bandwidth of 10 km and then utilizes an inverse-distance weighting for all points beyond this threshold. We chose this relatively small bandwidth to detect clusters at a relatively fine spatial level, while the inverse distance weighting beyond the 10-km bound prevents unjustified truncation of the appropriate neighborhood for detecting clustering. Nevertheless, the qualitative picture of clustering is robust over a range of alternative weighting strategies and bandwidth selections. The

calculation of G_i^* statistics was conducted using the hot spot analysis function in ESRI ArcToolbox.

Question 2: How was red king crab bycatch affected by the closures?—To address this question, we analyze summary statistics of the mean number of RKC caught per hour of trawling for the years before and after the closures, as well as the proportion of hauls for which recorded bycatch of crab is zero. The latter statistic is essential, since the catch data for RKC are highly zero-inflated; the majority of hauls yield little or no bycatch while a small share of “disaster” tows significantly influence the mean (Dew and Austring 2007).

In conducting these comparisons, we consider two subsets of our sample. First, we analyze the mean RKC per hour and the proportion of zero hauls over the entire available fishing grounds in each year (including the closure areas before 1995) and compare their trends to fishery-independent estimates of biomass. This provides a sense of the overall trend in RKC bycatch before and after the closures and suggests whether the observed trend is potentially explained by contemporaneous trends in crab abundance rather than closure implementation. Second, we analyze the same measures using only data from outside the closure areas. If the trends noted in the overall data set are absent in the non-closure data then that provides strong evidence that the observed changes are likely the product of displacement from the closure areas and are therefore attributable to the closure measures themselves.

As a more formal comparison of the distributions of RKC bycatch rates before and after the closure we partition the data into zero and non-zero observations and analyze each partition separately. To determine whether the proportion of observations with zero bycatch is higher after the closures, we utilize a two-sample test for the equality of proportions. We then utilize a Wilcoxon rank-sum test (Lehmann and D’Abrera 1998) on the non-zero RKC bycatch data to determine whether the pre-closure and post-closure distributions of crab bycatch encounters are equivalent.

Question 3: How did the implementation of closures alter the targeting behavior of fishermen?—In addition to the spatial displacement of effort addressed in the previous two questions, there may also be behavioral effects of the closures on fishermen’s targeting behavior. In particular, the closure may have displaced effort away from the primary target species, rock sole, to greater targeting of Pacific cod. We provide preliminary evidence to examine this question by examining the quantiles and central tendency of the distributions of rock sole and cod catch rates before and after the 1995 closures. To control for the effects of effort displacement from the closed areas to the remaining fishing grounds, we also examine the pre/post distributions of catch rates in the non-closure grounds. We use the Wilcoxon rank-sum test to formally test whether the catch rate distributions of each target species are comparable before and after the closures, or if one rank-dominates

the other. A significant upward shift in the catch rate distribution of cod that is not met by a similar increase for rock sole and that remains even after limiting the sample to non-closure areas may be indicative of greater relative targeting of cod triggered by the closure.

Changes in catch rates of rock sole and cod across seasons could arise due to shifting of effort across vessels with different targeting strategies or due to the redistribution of effort within seasons with intra-seasonal variation in catchability. Controlling for these possible confounding factors is necessary to avoid spurious identification of targeting changes. Our objective is to insulate our estimates of the interannual variation in expected catch rates of rock sole and cod from these confounding factors. We therefore posit the following baseline model for the conditional expectation of haul-level catch of rock sole and cod:

$$E[\text{Catch} | i, y, w, \text{Duration}] = \text{Duration} \times \exp(\alpha_i + \beta_w + \gamma_y) \quad (1)$$

where i , y , and w are indices for vessel, year, and week of season, respectively, Catch is catch of either target species in kilograms and Duration is the duration of a haul (time from deployment to retrieval of trawl gear) in hours. Our specification includes factor variables for vessel, week, and year, with the associated parameters indicated by α_i , β_w , and γ_y respectively. We treat each factor variable as a fixed effect, estimating it directly rather than using a random effects (i.e., multilevel) approach. This approach is justified given the available degrees of freedom in the data and is taken to avoid potential biases from the stronger assumptions underlying the random effects (i.e., multilevel) approach.

We estimate the model twice for each of the two target species. The first estimation uses the entire sample, while the second uses only the observations outside of the closures. Our primary interest is in the estimates of γ_y . To compare these estimates across years, they are exponentiated to provide the incidence rate of CPUE in each year relative to the base year, an estimate of the proportional scaling of CPUE controlling for vessel heterogeneity and week of participation. We set the base year to 1994, the last pre-closure year, by excluding this level of the factor variable from the specification (i.e., by setting $\gamma_{1994} = 0$).

Our primary concern in estimation is to achieve consistent estimates and robust inference for the γ_y factor variables. Assuming that the conditional expectation of catch rates is properly specified, consistent parameter estimates can be obtained for a GLM with a log-link function by maximizing the Poisson log-likelihood (Gourieroux et al. 1984b, McCullagh and Nelder 1998), using the duration of fishing as an exposure variable. Due to the membership of the Poisson in the linear exponential family of distributions, this result holds even if the Poisson is a poor description of the underlying distribution of the data (Gourieroux et

al. 1984a). To make our inference on the factor variables robust to failure of the Poisson likelihood as well as heteroscedasticity and possible spatial and temporal correlation, we utilize cluster-robust standard errors with clusters defined over unique combinations of vessels, weeks and years (Cameron and Trivedi 2005, Fitzmaurice et al. 2011). We estimate our models using the “Poisson” command with the cluster variance option in Stata software. This estimation procedure is equivalent to adopting a generalized estimating equations (GEE) approach to a Poisson GLM with a log-link function (Liang and Zeger 1986, Zeger and Liang 1986) in which the working variance for the GEE estimator takes the form prescribed by the Poisson distribution and where clustering in the robust (i.e., sandwich) errors is defined over combinations of vessels, weeks and years (Hardin and Hilbe 2003, Koper and Manseau 2009).

While analysis of changes in catch rates is suggestive, the fishery has substantial discard rates so that targeting cannot be inferred from catch composition alone. The observer program does not provide reliable estimates of discards for this time horizon. However, we can investigate changes in targeting by observing changes in weekly retained catch of each species estimated from the composition of each vessel’s weekly production. To examine how this retained catch varied after the closures in 1995, and to control for vessel-level differences in production and possible intra-seasonal trends, we estimate a linear regression explaining each vessel’s weekly retained mass of cod or rock sole per hour of fishing (Mg/h; 1 Mg = 1 metric ton):

$$\text{Retention}_{iwy} = \alpha_i + \beta_w + \gamma \text{PostClosure} + \varepsilon_{iwy} \quad (2)$$

where the α_i are vessel factors, β_w are weekly factors, and PostClosure is a variable that equals 1 when the closures are in place and is zero otherwise; γ is the ceteris paribus change in the retained catch per hour of fishing for a species after the 1995 closures. We estimate this regression using weighted least squares with the recorded hours of fishing by a vessel as the weighting variable (therefore weighting active vessels more highly) and utilize cluster-robust standard errors with clusters defined over vessel-years to account for heteroscedasticity and serial correlation in weekly production within years (Cameron and Trivedi 2005, Fitzmaurice et al. 2011).

Finding significant changes in catch and retention practices for cod and rock sole after 1995 is not sufficient to infer that the closures caused a shift in targeting behavior. One potential confounder may be changes in the absolute and relative biomass of cod and rock sole. To investigate this possibility, we compare our estimates of the out-of-closure trends in rock sole and cod catch from Eq. 1 to fishery-independent estimates of the biomass of each species. If the trends in biomass estimates are inconsistent in direction or magnitude with the trends in estimated catch rates outside the closures then this argues against biomass trends as an

explanation. To conduct a similar test using retention data, we modify Eq. 2 to include annual factor variables in place of the post-closure dummy variable:

$$\text{Retention}_{iwy} = \alpha_i + \beta_w + \gamma_y + \varepsilon_{iwy}. \quad (3)$$

By comparing the estimated pattern of γ_y to the year to year changes in estimated biomass, we can observe whether there is a strong coupling between biomass and the trends in retention of target species.

Question 4: What were the direct and indirect effects of the closures on halibut bycatch?—To investigate the direct effects of the closures on halibut bycatch rates due to displacement, we estimate exponential regressions of halibut catch that are identical to those previously described for rock sole and cod in Eq. 1. As before, these regressions isolate interannual changes in relative CPUE for halibut from changes due to the relative participation of different vessels and intra-seasonal patterns in bycatch. By comparing the proportional increases in halibut CPUE after the closures that are implied by estimates from the full sample vs. those from the out-of-closure sample, we are able to infer how much of any increase in halibut CPUE is directly attributable to displacement of fishing effort from the closure areas. By comparing the changes in CPUE over time to fishery-independent estimates of halibut biomass we can also assess to what degree any change in halibut bycatch outside of the closures is explained by trends in halibut biomass rather than alterations in fishing behavior.

Using the difference in CPUE changes between the full and out-of-closure sample to infer the effect of closure-induced displacement requires that the proportional difference between halibut bycatch rates in and out of the closure in 1994 is stable from 1995 onward. Since there is no fishing in the closed areas in the immediate aftermath of the closures, this assumption is not directly testable. However, we utilize several strategies to indirectly assess its plausibility. First, we estimate the same exponential regression specification as above using only the pre-closure (1992–1994) observer data but also interact each annual factor variable with a variable (Closure) that equals one when fishing occurs in a closure and is zero otherwise. The resulting specification of the conditional mean is

$$E[\text{Halibut} | i, y, w, \text{Duration}, \text{Closure}] \\ = \text{Duration} \times \exp(\alpha_i + \beta_w + \gamma_y + \delta_y \text{Closure}) \quad (4)$$

where δ_y is the coefficient attached to the interaction of an annual factor variable (indexed by y) with the closure indicator. Stability in the δ_y estimates relative to 1994 would suggest a temporally consistent wedge in halibut CPUE between fishing grounds in and out of the closure. We estimate Eq. 4 using the same GEE procedures described for Eq. 1.

As a second indirect test, we utilize the fortunate (for our purposes) coincidence that fishing was allowed in 1999 in the Special Savings Area of the RKCSA, an area

along the bottom of the closure that regularly saw roughly half the fishing in the RKCSA, to provide a glimpse of halibut bycatch rates within the closure after its establishment. We augment the 1992–1994 data set from estimating Eq. 4 to include the 1999 data as well and limit the within-closure data to only the hauls occurring inside the Special Savings Area, and then estimate the same specification. Comparing estimates of the proportional change in CPUE from fishing inside the closure in 1994 relative to 1999 can provide further evidence for or against the stability of the halibut bycatch gradient since 1994.

As a third test, we proxy for the lack of any fishing within the closures after their implementation by examining the pattern of halibut CPUE in the area to the immediate south of the RKCSA. This area is a narrow rectangular strip extending 28 km from the RKCSA's southern border and extending over its entire 125 km width. Given its spatial proximity, this area may exhibit some similarities to the once heavily-fished adjoining areas inside the RKCSA. Once again, we estimate the following specification for the conditional mean of halibut bycatch using the GEE procedures described for Eq. 1

$$E[\text{Halibut} | i, y, w, \text{Duration}, \text{Closure}] \\ = \text{Duration} \times \exp(\alpha_i + \beta_w + \gamma_y + \delta_y \text{Closure}). \quad (5)$$

However, two important changes are made to the sample and definition of the variables. First, we now limit the sample to only the data outside of the closure areas and define Closure = 1 when fishing occurs in the area to the immediate south of the RKCSA and = 0 otherwise. Second, since our objective is to examine the long-run stability of relative halibut bycatch rates in this area, we expand the sample to include all data from 1992 to 1999. Estimates of the δ_y factors for 1995 onward that are stable in magnitude and indistinguishable from δ_{1994} support the assertion that the wedge between halibut bycatch rates inside and outside the closure in 1994 would have persisted in the absence of the closures.

To explore the indirect impact of the closures on halibut bycatch through their effects on the relative targeting of rock sole and cod, we first investigate whether there is evidence of any significant correlation between cod and halibut CPUE. Preliminary evidence suggested that halibut and cod are log-linear in relationship. We therefore take natural logarithms of halibut and cod catch rates for each haul and calculate their correlation for the 78% of our sample for which halibut bycatch rates are non-zero. We separately estimate a logit model to explore any link between cod catch rates and the probability of zero halibut bycatch. We also examine whether the linkage between cod and halibut catch rates is driven indirectly through a correlation between cod and rock sole catch rates by examining the correlation between the logarithms of cod and rock sole catch rates and also investigating whether

the partial correlation between cod and halibut is significantly decreased by first controlling for rock sole.

To examine the effects of targeting in a more rigorous way that controls for potentially correlated factors, we supplement the halibut CPUE model on the non-closure data with regressors for the catch (per hour) of cod and rock sole, estimating the following conditional mean:

$$E[\text{Halibut} | i, y, w, \text{Duration}, \text{Cod}, \text{Rsole}] \\ = \text{Duration} \times \exp(\alpha_i + \beta_w + \gamma_y + \delta_{\text{cod}} \text{Cod} \\ + \delta_{\text{rsole}} \text{Rsole}). \quad (6)$$

Once again, we estimate this specification using the GEE methods described for Eq. 1. By examining the relative magnitude of δ_{cod} and δ_{rsole} , we can assess the relative impact of shifting the composition of a metric ton of catch between rock sole and cod.

Since both rock sole and cod catch rates varied in the wake of the closures and interact in their influence on halibut bycatch rates through the multiplicative nature of the exponential regression specification, we cannot assess the effects of shifts in targeting based upon parameter estimates alone. Instead, we utilize the parameter estimates from the regression model as the basis for a simulation to isolate the ceteris paribus effects of shifts in targeting behavior. To do this, we first assume that all variables in the post-closure data aside from target catch rates in our regression (i.e., vessel identity and timing of fishing) remain constant in the simulation. We further assume that the realizations of stochastic errors (the variables that resolve the gap between the observed catch rate and the predictions from our estimated model) in the exponential regression are also fixed at their implicit values in the 1995–1997 data and are multiplicative to the exponential specification of the conditional mean of catch. Under these assumptions, the ceteris paribus proportional change in halibut CPUE for any simulated changes in the catch rates (per hour) of cod (ΔCod) and/or rock sole (ΔRsole) is

$$\text{Prop}\Delta\text{CPUE} = \exp(\delta_{\text{cod}}\Delta\text{Cod} + \delta_{\text{rsole}}\Delta\text{Rsole}). \quad (7)$$

Our simulation takes the post-closure data as given and entertains the question of how halibut bycatch outside the closure areas changed relative to a counterfactual baseline in which rock sole and cod catch rates are drawn from their 1994 empirical joint density. The simulation proceeds in four steps. First, for each 1995–1997 observation data point, we draw one observation with replacement from the 1994 data on cod and rock sole catch rates in non-closure areas. Second, we calculate the implied changes in the target species catch rates for each observation (ΔCod and $\Delta\text{RockSole}$), treating the 1994 draw as the baseline scenario. Third, we use Eq. 7 to calculate the proportional change in CPUE for each post-closure observation. For each draw, we find the ceteris paribus effects of changes in cod and rock sole by holding the other target species at

its realized post-closure level. We also calculate a measure of $\text{Prop}\Delta\text{CPUE}$ in which both species vary simultaneously. Fourth, we take the annual means between 1995–1997 of the three simulated proportional changes to find the average proportional changes in CPUE that arise for one realization of the simulation. We then repeat these four steps over 1000 simulations and take the overall mean of these simulations for our final predictions of the expected effect in each year. All simulations were conducted using Stata 11 software.

RESULTS

Question 1: How did fishing effort redistribute in the wake of the closures?—Between 1992 and 1994, 47% of trawls in the fishery occurred in the RKCSA, with only 8% occurring in the PHCA. Fig. 1 shows that the redistribution of effort after the closures was primarily concentrated in three zones: along the southern border of the RKCSA, to the southwest of the RKCSA, and in the southern end of the grounds.

Fig. 2 shows patterns of clustering both before and after the closures using the G_i^* statistic. Not surprisingly, given the stated objectives of the RKCSA, red king crab had substantial hotspots within its boundary with little pre-closure evidence of the hotspot to the southeast of the RKCSA that is evident from 1995 onward. However, there was relatively little fishing in this region before 1995. Prior to the closure, halibut exhibits a similar pattern to 1995–1997 with some evidence of an area of persistently high bycatch in the western RKCSA. Nevertheless, there is little overlap in the hot spots of these two bycatch species. The closures also contained the primary “hot spots” for rock sole fishing, not surprising given the concentration of fishing effort in these areas.

A comparison of the bottom panel of Fig. 2 with the redistribution of effort shown in Fig. 1 suggests that the redistribution of effort closely aligns with aggregations of rock sole and cod in these areas. Overall, fishermen increased effort in relatively halibut- and cod-rich waters once the RKC and rock sole-rich areas inside the closures were placed off-limits.

Question 2: How was red king crab bycatch affected by the closures?—Table 1 shows the mean RKC bycatch rates (number of crabs per hour of hauling time), estimates of RKC biomass, as well as the proportion of hauls for which recorded crab bycatch is zero. The mean bycatch rate fell precipitously from 13 crab per hour for 1992–1994 to 3.3 from 1995–1997 without any reduction in surveyed crab biomass that would help explain this decline. Similarly, the proportion of zero-RKC hauls increased dramatically in the post-closure years.

Table 1 also includes bycatch rates and proportions of zeros for trawls outside of the closure area. These statistics reveal, with the exception of the unusually high bycatch rates in 1994, that the patterns of crab bycatch in the years after the closures are comparable to those before. The hypothesis that the proportion of hauls with

zero crab are equal between 1992–1994 and 1995–1997 cannot be rejected ($Z = -1.44$, $P = 0.15$). A Wilcoxon rank-sum test comparing the distribution of crab encounters from 1992–1994 to 1995–1997 yields a Z statistic of 0.028 ($P = 0.977$). The distribution of RKC bycatch outside the closures is therefore quite stable through time. This invariance, paired with the lack of an explanatory trend in RKC biomass, strongly supports the hypothesis that the dramatic reductions in RKC bycatch are attributable to the relocation of effort from the closures. Thus, the closures, and in particular the RKCSA, were highly successful in reducing the impact of fishing on red king crab.

Question 3: How did the implementation of closures alter the targeting behavior of fishermen?—While the foregoing analysis has emphasized the spatial displacement of effort from the closures, there was also a significant but less obvious behavioral effect of the closures on fishermen’s targeting behavior. Table 2 shows the distributions of rock sole and cod catch rates from the observer data sample. There is strong evidence of a distributional shift for both species after 1995, with a decrease in rock sole catch rates and a strong increase in cod catch rates at the mean and for all quantiles. Once we control for the direct effect of displacement by examining only the data from grounds outside the closures, the downward shift is largely swept away for rock sole (with the Wilcoxon rank-sum test insignificant) while a large and significant upward shift in the distribution of cod catch rates persists.

The incidence rates associated with the estimates of the annual factor variables from Eq. 1 are presented in Fig. 3 for both rock sole and cod. To examine the importance of controlling for vessel and weekly fixed effects, we also calculate the mean CPUE in each year of the raw data for both species and standardize these estimates by their 1994 means. Fig. 3 confirms the findings of the summary statistics, revealing a strong post-closure erosion in rock sole CPUE offset by a substantial increase for cod. Comparing the estimates from the entire data set to those estimated using only data outside the closures, we fail to identify the 1995–1997 reduction in rock sole CPUE. This suggests that the reduction in rock sole CPUE was strongly driven by displacement to less favorable grounds. However, the upward shift in cod CPUE is robust to this constrained sample, which signifies a marked change in either the abundance of cod or the tendency to target it.

The estimates of γ from Eq. 2 are the ceteris paribus change in the retained catch per hour of fishing for a species after the 1995 closures. The estimated effect for rock sole is -0.119 Mg/h ($Z = -1.57$, $P = 0.116$). Compared to a (fishing hours weighted) pre-closure mean of 1.29 Mg/h, this effect is small and insignificant, indicating that retention of rock sole for production was relatively steady post-closure. However, the estimated effect for cod is 0.51 Mg/h ($Z = 6.41$, $P = 1.5 \times 10^{-10}$). Given its pre-closure mean retention rate of 0.21 Mg/h,

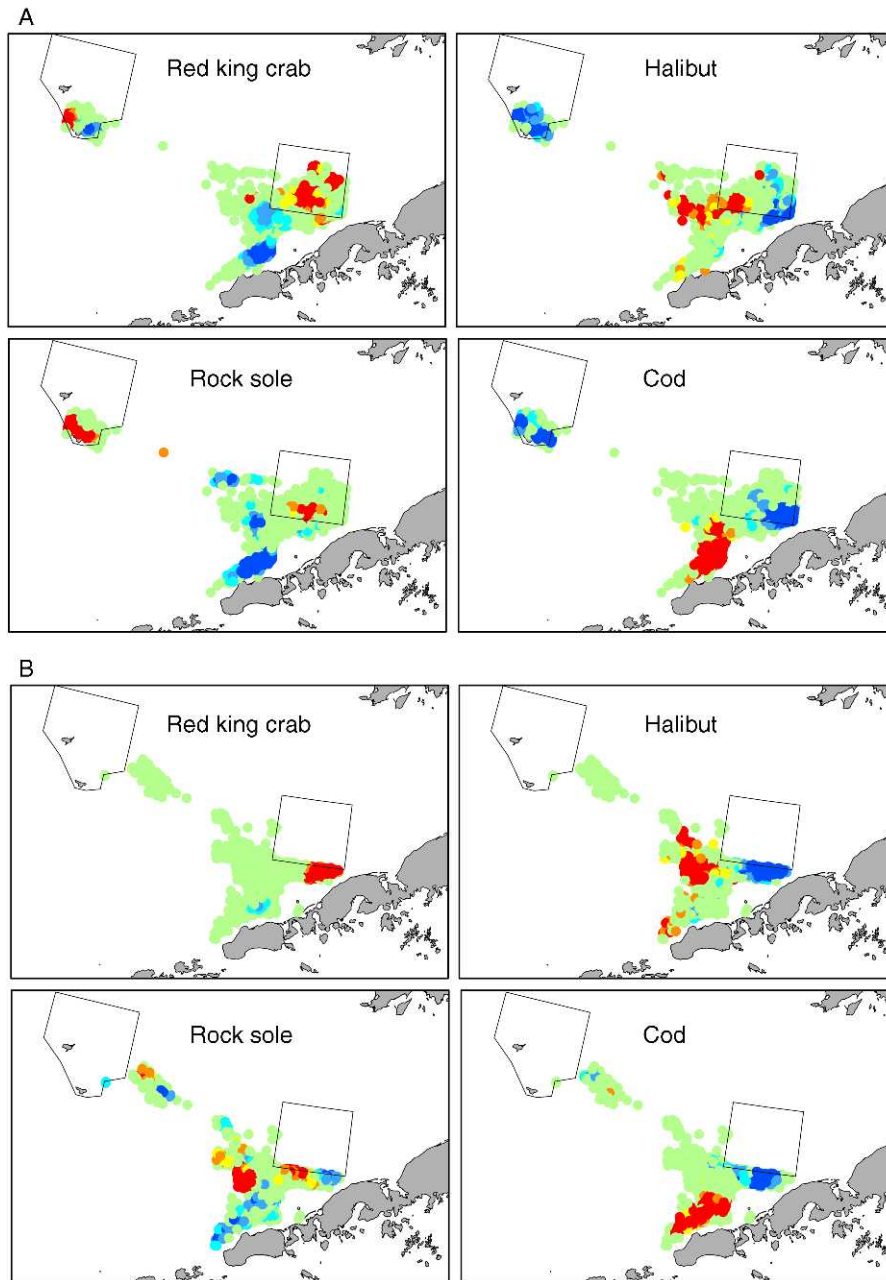


FIG. 2. Z scores from G_i^* statistics for four species indicating likely areas of clustering of high and low catch rates in (A) 1992–1994 and (B) 1995–1997. Observations from yellow to red indicate likely “hot spots” (yellow, $1.65 < z < 1.96$; orange, $1.96 < z < 2.58$; red, $z > 2.58$) while observations in blue indicate potential “cold spots” with clustering of low catch rate areas (light blue, $-1.96 < z < -1.65$; medium blue, $-2.58 < z < -1.96$; dark blue, $z < -2.58$). Areas in green indicate no clustering of high or low catch rates ($-1.65 < z < 1.65$).

this implies an increase of over 140% in the retention of cod in a typical fishing hour. The increases in cod catch after the closures were therefore met by roughly proportional increases in retention, supporting the notion of increased cod targeting from 1995 onward.

Fig. 4 shows fishery-independent estimates of the Eastern Bering Sea biomass of cod and the female spawning biomass of rock sole. We focus on female

spawning biomass for rock sole, since there is very little retention or targeting of male or immature female rock sole in the winter fishery. Note that the estimated cod biomass peaks in 1994, the year *before* the implementation of the closures, and declines steadily thereafter before returning to 1992–1993 levels by 1997. The three-year increase of cod from its 1992–1993 levels occurs a year too early to explain the shift toward cod

TABLE 1. Annual comparisons of mean red king crab (RKC) bycatch rates and proportion of hauls with zero bycatch (Prop. = 0) from observer data relative to estimates of biomass.

Year	All areas		Non-closure areas		RKC biomass (millions of Mg)		
	Bycatch rate (no./h)	Prop. = 0	Bycatch rate (no./h)	Prop. = 0	Male	Female	Total
1992	6.10	0.86	1.49	0.95	18.30	15.60	33.90
1993	15.20	0.74	2.01	0.92	25.70	21.70	47.40
1994	16.97	0.71	7.70	0.82	20.40	13.50	33.90
1995	3.95	0.91	3.95	0.91	24.60	15.30	39.90
1996	1.87	0.92	1.87	0.92	26.70	26.60	53.30
1997	3.30	0.95	3.30	0.95	48.50	28.30	76.80

Note: Means are weighted estimates calculated from haul-level data using the duration of haul as the weight. Note that 1 Mg = 1 metric ton.

production. Estimates of the yearly factor variables from Eq. 3 are reported in Table 3 and show no evidence of a pattern of production in 1995–1997 that mirrors the observable trend in cod biomass for this period. In fact, the mean production of cod in 1997 exceeded that in 1995 despite an estimated biomass of only half the 1995 level. There is also no evidence of an uptick in production in 1994 to coincide with the biomass spike estimated in the trawl survey in the summer of that year. Instead the increase in production occurs in 1995, after the closures were in effect. Finally, rock sole biomass increased substantially over the sample period, so an increase in the relative abundance of cod over this horizon is not a viable explanation for the dramatic observed changes in catch and production.

Fig. 4 also reports the mean ex-vessel prices for the dominant product types for each species. Rock sole prices were strong (until a crash in 1997) even as cod prices declined slowly through time. There is therefore no evidence of a market-driven rationale for the switch toward more intense cod catch and production. The biomass and product-market hypotheses do not adequately explain the shift toward cod harvest and production in 1995, leaving the closures themselves as

the most probable cause. Driven from their prime rock sole fishing grounds, fishermen adapted to lower rock sole catch rates by seeking out and targeting increased “bycatch” of cod and increasingly retaining it for production.

Question 4: What were the direct and indirect effects of the closures on halibut bycatch?—Table 4 shows the distributions of bycatch rates of halibut before and after the implementation of the closures. There is a large upward shift in both the mean and overall distribution of halibut after the closures went into effect as reflected by the significant Wilcoxon test ($Z = -4.37, P = 1.2 \times 10^{-5}$). There is also a significant reduction in the proportion of hauls with zero halibut ($Z = 4.86, P = 1.2 \times 10^{-6}$). When this comparison is limited to fishing in the non-closure areas, the post-closure increase in the mean shrinks from over 36 to 21 kg/h and the Wilcoxon test remains significant ($Z = -2.21, P = 0.027$), so that average halibut bycatch increased even after accounting for displacement from the closure. This change was not uniform across the quantiles, however. The proportion of zero halibut hauls actually *increased* slightly after the closures ($Z = -1.97, P = 0.049$), and the lower quantiles (including the median) actually fell somewhat. These

TABLE 2. Comparisons of the distributions of target species before and after the closures between the entire sample and the subsample occurring outside the closure areas.

Timing	CPUE (Mg/h)						Z	N
	Mean	25th	Median	75th	95th	Maximum		
All areas								
Rock sole								
Before	4.03	1.98	3.41	5.11	9.55	53.00	10.86	2304
After	3.33	1.28	2.55	4.55	9.32	63.03		
Cod							-27.9	2304
Before	0.72	0.24	0.50	0.92	2.03	9.40		
After	1.54	0.63	1.12	1.87	4.20	17.06		2210
Non-closure area								
Rock sole							1.28	1023
Before	3.16	1.44	2.71	4.25	7.10	27.20		
After	3.33	1.28	2.55	4.55	9.32	63.03		2210
Cod							-13.72	1023
Before	0.97	0.35	0.70	1.23	3.06	9.40		
After	1.54	0.63	1.12	1.87	4.20	17.06		2210

Notes: The mean and quantiles are calculated from haul-level data using the duration of tow as a weight on each observation. Z is from the Wilcoxon rank-sum test comparing before and after; N is the number of observations.

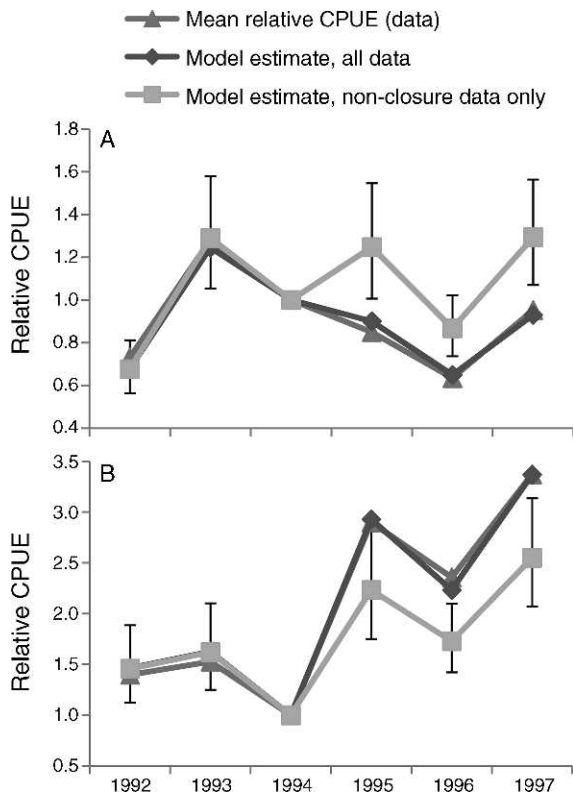


FIG. 3. Graphs of observed (triangles) and estimated proportional catch per unit effort (CPUE) relative to 1994 (i.e., 1994 CPUE = 1) for (A) rock sole and (B) cod. Estimates are derived from robust count regressions with vessel and weekly controls using all data (diamonds) and only data from outside the closures (squares). The 95% confidence intervals for non-closure estimates are included.

changes were more than offset, however, by increases in the upper quantiles of halibut CPUE. The post-1995 increase in the mean bycatch rate is wholly attributable to a “fattening” of the upper tail of the bycatch distribution.

Fig. 5 shows the measures of relative halibut CPUE resulting from the exponentiated estimates of annual factor variables from Eq. 1. These reveal a strong uptick in CPUE in post-closure years with rates between 96–143% of 1994 levels. Surveyed halibut biomass for this region over the same period exhibited little trend and is not consistent with this increase. Limiting the sample to data from outside the closures dampens the increase somewhat, with halibut bycatch in non-closure areas increasing by 52–92% after the closure relative to 1994. This significant dampening of the post-closure increases in halibut CPUE suggests that a substantial share of the post-closure increase in halibut bycatch is attributable to displacement from the closed areas.

This inference relies on an assumption that the wedge between halibut bycatch rates in and out of the closures in 1994 is stable for the years immediately after the closure. In other words, this assumes that opening the

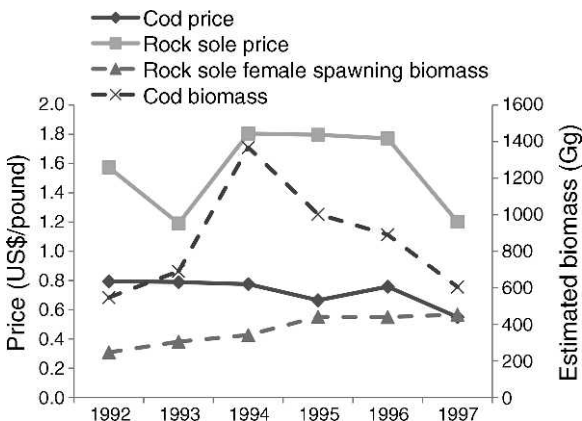


FIG. 4. Price and biomass trends for cod and rock sole in the Eastern Bering Sea (1 pound = 0.45 kg; 1Gg = 1000 metric tons). Biomass estimates are based on fisheries independent surveys.

closures would lead to a predictable reduction in halibut bycatch. The estimates from Eq. 4 for the incidence rates associated with fishing in a closure (relative to the baseline of fishing outside the closures) in 1992–1994 are reported in the first column of Table 5. Fishing in and out of closures yielded identical expected halibut bycatch in 1992 and 1993. However, in 1994, fishing in the closure yielded only 60% of the CPUE of halibut as fishing outside, suggesting that this year either marked the beginning of a new pattern or is a misleading aberration.

To shed light on these possibilities we examine the evidence provided by the opening of the Special Savings Area (SSA) of the RKCSA in 1999. As discussed in the methods section, we estimate a modified version of Eq. 4 by augmenting the 1992–1994 data set from the previous regression to include the 1999 data as well and limit the within-closure data to those hauls inside the SSA of the RKCSA. The estimates of the relative CPUE in each year from fishing in the SSA are reported in the second column of Table 5. The pattern of CPUE for 1992–1994 corresponds closely to the estimates in the first column (which are estimated using all observations within the closures), suggesting the SSA is a reasonable proxy for not having observed fishing in the broader RKCSA over this period. Most importantly, the estimates for 1994 and 1999 are both statistically significant, indistinguish-

TABLE 3. Estimates of annual fixed effects from cod retention (Mg/h) regression (Eq. 3) (base year = 1994).

Year	Estimate	95% CI
1992	0.255 (2.45)	0.047, 0.463
1993	-0.074 (-0.77)	-0.263, 0.116
1995	0.586 (6.12)	0.394, 0.777
1996	0.496 (6.25)	0.338, 0.655
1997	0.750 (7.22)	0.543, 0.958

Note: Z statistics from cluster-robust standard errors are in parentheses.

TABLE 4. Comparisons of the distributions of halibut bycatch rates before and after the closures between the entire sample and the subsample occurring outside the closure areas.

Halibut (kg/h)	Prop. = 0	Bycatch rate (kg/h)						Z		N
		Mean	25th	Median	75th	95th	Max	Wilcoxon rank-sum test	Two-sample proportion test	
All areas										
Before	0.26	81.54	0	49.24	113.63	290.83	1141.43	-4.38	4.86	2304
After	0.20	118.09	10.49	56.19	151.98	449.69	1395.52			
Non-closure area										
Before	0.17	96.82	21.58	65.00	133.24	317.33	1141.43	-2.21	-1.97	1023
After	0.20	118.09	10.49	56.19	151.98	449.69	1395.52			

Notes: Means, proportions, and quantiles are calculated from haul-level data using the duration of tow as a weight on each observation. The Wilcoxon rank-sum test evaluates the null hypothesis that neither the pre- nor post-closure distribution of positive bycatch observations dominates the other in rank order. It is therefore conducted using only observations for which halibut bycatch is positive. The two-sample proportion test tests the null hypothesis that the proportion of zero bycatch hauls is identical before and after the closures.

able in magnitude, and suggest a substantial reduction in mean halibut CPUE from fishing inside the SSA in both years.

As a further indirect test, we examine the bycatch gradient between the area to the immediate south of the RKCSA and the remaining open areas (Eq. 5). The estimates for the relative CPUE in the RKCSA-adjacent open areas compared to other open grounds are provided in Fig. 6. The pattern and magnitude of these estimates coincides closely with previously noted patterns of relative halibut CPUE between the RKCSA and all other open areas; bycatch rates across the zones are indistinguishable in 1992 and 1993 and rates inside the area to the immediate south of the RKCSA are roughly 56% of the bycatch rates in other open areas in both 1994 and 1999. Furthermore, bycatch rates are consistently lower in this boundary area from 1994 to 1999 than in other non-closure grounds (with the exception of 1997, which is insignificantly different) and are also fairly stable in their estimated values. Altogether, evidence suggests that 1994 was not an aberrant year; halibut bycatch rates would likely have been significantly lower inside the RKCSA if fishing had been allowed there from 1995–1997. This further buttresses our contention that substantial increases in halibut bycatch rates are directly attributable to displacement from the closure areas.

While effort displacement is an important aspect of the post-1995 increase in halibut bycatch, Fig. 5 shows that a significant increase relative to 1994 persists even after controlling for this effect. Some of this increase can be explained by the shift of targeting toward cod in the wake of the closures. Fig. 7 plots logged halibut and cod catch rates, revealing a significant positive correlation ($\rho = 0.23, P < 0.0001$). This relationship is not an artifact of correlation between the two target species as the log catch rates of cod and rock sole are negligibly correlated ($\rho = 0.02, P = 0.132$), and the partial correlation between cod and halibut is minimally affected by controlling for rock sole, falling from 0.23 to 0.21.

Estimates of Eq. 6, reported in Table 6, reveal that a 1 metric ton (Mg) increase in rock sole catch yields an increase in halibut bycatch of 5% while an equivalent increase in the catch of cod leads to a 10% increase. A Wald test (Cameron and Trivedi 2005) strongly rejects the hypothesis that cod and rock sole targeting have equivalent effects on halibut bycatch rates ($P < 0.01$). Fig. 5 shows the results of controlling for targeting on estimated halibut CPUE relative to 1994 levels. Notably, the effects on estimates are small in pre-closure years but are substantial afterward with the annual increases in CPUE compared to the baseline of 1994 declining from 52% to 30% in 1995, 53% to 43% in 1996 and 92% to

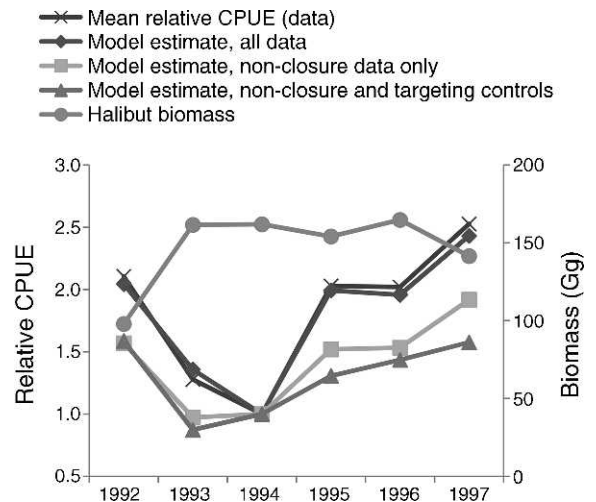


FIG. 5. Graph of observed (crosses) and estimated proportional CPUE relative to 1994 for halibut where effort is defined as a vessel-hour of trawl time. Estimates are derived from robust count regressions with vessel and weekly controls using all data (diamonds), only data from outside the closures (squares), and data from outside closure areas controlling for rock sole and cod catch rates (triangles). Estimates of halibut biomass are included for comparison (circles); 1 Gg = 1000 metric tons.

TABLE 5. Annual estimates of halibut CPUE (kg/h) inside closures relative to outside (i.e., incidence rates of δ_y from Eq. 4).

Year	All grounds	Non-closure area + RKCSSA
1992	1.029 (0.19)	1.001 (0.00)
1993	1.106 (0.63)	0.968 (-0.16)
1994	0.607 (-3.12)	0.675 (-1.87)
1999		0.656 (-2.43)

Notes: Z statistics from cluster-robust standard errors are in parentheses. The first column reports estimates for the full sample of 1992–1994 data. The second column reports estimates for a sample in which closure fishing is limited to the special savings area (RKCSSA) of the red king crab savings area (RKCSA) closure, and 1999 data are included. See *Models: Question 3: How did the implementation of closures alter the targeting behavior of fishermen?* for an explanation of incidence rates.

58% in 1997 after controlling for the change in catch composition.

Our simulation of the 1995–1997 halibut CPUEs under the counterfactual that either cod or rock sole catch rates (or both) were drawn from their 1994 empirical distribution shows that the percentage changes in halibut CPUE attributable to the increases in cod targeting (holding rock sole constant at realized post-closure values) are 12% (1995), 8% (1996), and 14% (1997) if all other post-closure factors remained as they are in the data. The equivalent ceteris paribus measures for rock sole are 2%, -3%, and 8%, respectively. When rock sole and cod are varied simultaneously, the simulations show that targeting changes account for overall increases in out-of-closure halibut CPUE of 14%,

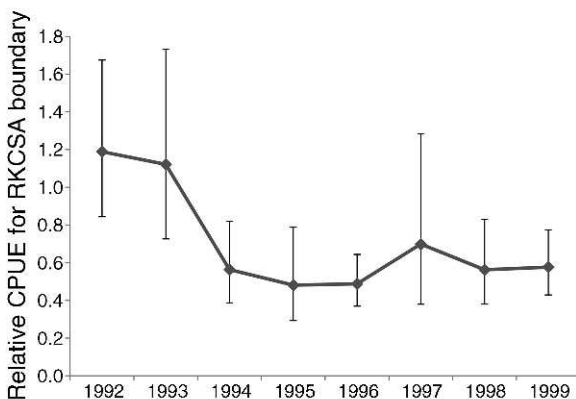


FIG. 6. Halibut CPUE (incidence rates of δ_y from Eq. 5; where δ_y is the coefficient attached to the interaction of an annual factor variable, indexed by y , with the closure indicator) of the area to the immediate south of the red king crab savings area (RKCSSA) relative to other areas outside the closures. The 95% confidence intervals for incidence rates are calculated by exponentiating the bounds of the asymptotic confidence intervals for the coefficient estimates. See *Models: Question 3: How did the implementation of closures alter the targeting behavior of fishermen?* for an explanation of incidence rates.

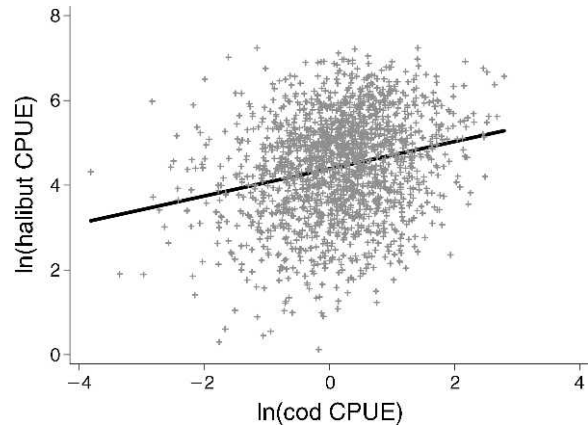


FIG. 7. Scatterplot and fitted trend of the natural log of halibut CPUE against the natural log of cod CPUE.

5%, and 26% respectively. This demonstrates the significant influence of targeting decisions on halibut bycatch and highlights the dominant role of the closure-induced targeting of cod in exacerbating this bycatch. It also provides an example of the importance of accounting for the secondary behavioral effects of closures, in addition to more easily anticipated primary effects. Secondary effects, such as targeting changes, are not guaranteed to exacerbate negative spillovers of spatial management for bycatch; however, they are more likely to do so in cases where desirable target species occur in close spatial association with bycatch species outside the closures and the closures restrict fishermen's ability to break this linkage by diversifying their fishing over space and time.

DISCUSSION

We utilize unusually high-quality data gathered before and after the implementation of permanent spatial closures to examine the closures' influence on commer-

TABLE 6. Estimates from exponential regression of halibut bycatch (kg) on year, week, and vessel factors and catch rates (no./h) of rock sole and cod using non-closure data only (Eq. 6).

Parameter	Coefficient	Incidence rate	Incidence rate 95% CI
1992	0.462 (2.30)	1.587	1.072, 2.351
1993	-0.136 (-0.59)	0.873	0.556, 1.370
1995	0.267 (1.29)	1.306	0.870, 1.961
1996	0.361 (2.00)	1.435	1.008, 2.042
1997	0.455 (2.30)	1.575	1.076, 2.307
Rock sole	0.052 (6.01)	1.053	1.036, 1.071
Cod	0.096 (6.91)	1.100	1.071, 1.131

Notes: Z statistics, derived from cluster-robust standard errors, are reported in parentheses. The 95% confidence intervals for incidence rates are calculated by exponentiating the bounds of the asymptotic confidence intervals for the coefficient estimates. See *Models: Question 3: How did the implementation of closures alter the targeting behavior of fishermen?* for an explanation of incidence rates.

cial fishermen's targeting behavior, catch and multispecies bycatch. We believe the Bering Sea fishery shares a number of important features with a wide array of multispecies fisheries, giving our findings applicability well beyond the immediate context. The use of closures and catch quotas is a common management tool for the protection of "weak" stocks in a multispecies context. However, our case study suggests several risk factors that could make proposed closures vulnerable to behavioral adaptations. These include (1) patterns of association of target and bycatch species that are quite distinct within vs. outside the closures, creating the capacity for an adaptive response that looks quite different than that observed in the past; (2) the closures cover areas with significant pre-closure effort, providing fishermen with greatly diminished fishing opportunities for their usual target species; (3) fishermen use a gear that can be adapted to alternative fishing grounds or target stocks at minimal cost; and (4) there are established markets for alternative target species.

Our results sound a cautionary note for the design and use of spatial closures as a tool for ecosystem-based fisheries management. While they may generate unique benefits relative to alternative policies, time-area closures in multispecies environments can increase fishing-related pressure on target and bycatch species in both direct and more subtle ways. To the extent that an ecosystem approach to fisheries management entails the careful consideration of balancing fishing impacts across target species and habitat types, MPAs and other forms of spatial closures may be excessively blunt and unpredictable instruments for achieving the biological objectives of ecosystem-based management. When fishing opportunities are significantly altered as a result of these policies, managers should assume that unintended behavioral adjustments are likely and be ready to adapt to these adjustments as they emerge. While techniques for predicting the reallocation of fishing effort to novel management restrictions exist (e.g., Smith and Wilen 2003, Haynie and Layton 2010), uncertainty in these predictions remains, and patterns of catch/bycatch rates observed directly before the closure may differ from those observed afterward. Experience has demonstrated that fishermen are often surprisingly adaptive in the wake of new regulations (Wilen et al. 2002). Anticipating these adjustments may be difficult, however, leading to reactionary policies aimed at redressing the undesirable symptoms of these adaptations. Unfortunately, the piecemeal sequence of regulations that evolves from this game of "cat and mouse" may be ineffective on both biological and economic grounds.

One alternative management system is to subdivide the allowable catch of all relevant target and bycatch species to individual fishermen, effectively creating a system of multispecies individual tradable quotas (ITQs), so that fishermen face individualized incentives to avoid each species. Such "incentive-based" policies

(Grafton et al. 2006, Costello et al. 2008) have been adapted successfully in a handful of multispecies fisheries (Sanchirico et al. 2006, Branch and Hilborn 2008). They have the merit of steering fishermen toward practices that cost-effectively achieve the biological objectives reflected in the quotas rather than requiring managers to engage in potentially costly and error-prone forecasting of the biological and economic effects of policies directed at regulating fishing behavior or technology rather than directly targeting the desired outcomes. Incentive-based policies can motivate fishermen to engage in rapid in-season adaptive management, capitalizing on their often superior short-run knowledge of the available tradeoffs across species. The cost effectiveness of such policies, particularly if quota is tradable across fishermen, reduces the economic burden on the fishing industry, potentially reducing opposition to more extensive conservation measures.

Nevertheless, incentive-based policies alone are not a panacea for the difficult task of multispecies fishery management. They may present unique monitoring challenges, and some current incentive-based policies may not offer sufficient protection of vulnerable species, life-stages, or sensitive habitats relative to spatial closures. One approach is to devise policies that merge the strengths of incentive-based policies with some of the spatial control of closures. For instance, a limited subset of target and bycatch ITQs could be specifically allocated to spatial zones so that only vessels possessing these specially designated shares can fish in these areas and the shares cannot be used to fish elsewhere. By allowing fishermen to trade the rights to fish within these zones, the spatial concentration of fishing is limited while affording fishermen a "safety valve" to fish outside of the regular fishing grounds if profit opportunities or the lower bycatch of protected species warrant the higher price for the privilege of fishing in the protected area. Once deemed unrealistic to enforce, such policies are increasingly feasible due to advancements in technology (Smith and Wilen 2002). Economists have also suggested the possibility of developing habitat quotas that would allow vessels to trade access to different fishing habitats based on the degree of impact that different fishing gears may impose (Holland and Schnier 2006). These innovations notwithstanding, MPAs and other closure policies will continue to play a valuable role in the ecosystem-based management of fisheries. However, as a polar solution along a continuum of increasingly feasible hybrid policies, their use should be targeted toward cases where the unique benefits they afford justify the possible economic and biological repercussions associated with their use.

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