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[Article]

Developing Model-Based Indices of Lake Whitefish Abundance Using Commercial Fishery Catch and Effort Data in Lakes Huron, Michigan, and Superior

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Abstract.—Fishery catch per effort (CPE) is often used to assess relative fish abundance, and in many Great Lakes and other freshwater applications the CPE index is based on either an average or the ratio of summed aggregate catch to summed aggregate effort. In particular, assessments used to estimate the abundance of lake whitefish *Coregonus clupeaformis* and recommend harvest quotas in 1836 Treaty-ceded waters of Lakes Huron, Michigan, and Superior assume that commercial CPE from gill-net and trap-net fisheries is proportional to abundance; however, CPE may change due to factors other than abundance, leading to violations of this assumption. To account for sources of CPE variation that are not attributable to abundance, general linear mixed models (GLMMs) were developed for each management unit and least-squares means (LSMs) for each year were used as the index of abundance. The effect of using the GLMM method for standardization was evaluated by examining the temporal trends in the proportional difference between LSMs and CPE (i.e., aggregate catch divided by aggregate effort for each year). Of the random effects included in the final GLMM for the gill-net fishery, license holder accounted for the most variation. The fixed effect of boat size category on CPE depended on lake; on average, there was little difference between boat sizes in Lake Superior, whereas large boats had a lower CPE than medium and small boats in Lakes Michigan and Huron. On average, CPE was higher from October to December than during other months. The proportional difference between LSMs and CPE trended through time in some management units, suggesting the importance of adjusting fishery CPE for effects like boat size, season, and license holder. Factors that influence lake whitefish commercial fishery CPE are similar to factors important for marine commercial fisheries.

The lake whitefish *Coregonus clupeaformis* has supported a historically important fishery for Native American bands and a highly valued commercial fishery in the upper Great Lakes (Lakes Huron, Michigan, and Superior). In the late 1800s and early 1900s, lake whitefish were often the most highly valued commercial species and usually comprised the greatest proportion of total yield from each of the upper Great Lakes (Koelz 1926; Brown et al. 1999). Lake whitefish stocks collapsed in each of these lakes during the 1930s and 1940s due to overexploitation, predation by sea lampreys *Petromyzon marinus*, and pollution (Smiley 1882; Koelz 1926; Jensen 1976; Brown et al. 1999; Ebener and Reid 2005). During the 1960s through the 1980s, lake whitefish stocks rebounded in each of the lakes largely due to improved management of commercial harvest, sea lamprey control, pollution remediation, and the introduction of salmonines that reduced the abundance of the invasive alewife *Alosa pseudoharengus* and rainbow smelt *Osmerus mordax* (Ebener 1997; Mohr and Ebener 2005a). In the 1990s, lake whitefish once again became the main commercial species, particularly in Lake Huron, where the species constituted over 80% of the total commercial yield (Mohr and Ebener 2005b).

In 1979, the rights of Native American bands to fish in the Michigan waters of the upper Great Lakes as reserved in an 1836 Treaty were reaffirmed by U.S. federal courts. Since the reaffirmation of treaty fishing rights, periodic stock assessments have been conducted for stocks within spatially defined management units; the fishery data and harvest from within each management unit are treated as applying to a reproductively isolated stock (Figure 1; Ebener et al. 2005). Stock assessments are conducted and harvest recommendations based on the assessments are made annually for each individual management unit. Within each management unit, commercial fishery catch and effort data are reported on a 10 × 10-min statistical grid basis, which allows for some spatial resolution within management units.

Guidelines for the management of lake whitefish have been set according to the 2000 Consent Decree for 1836 Treaty-ceded waters. The 2000 Consent Decree created a Technical Fisheries Committee and a Modeling Subcommittee (MSC) to conduct stock assessments

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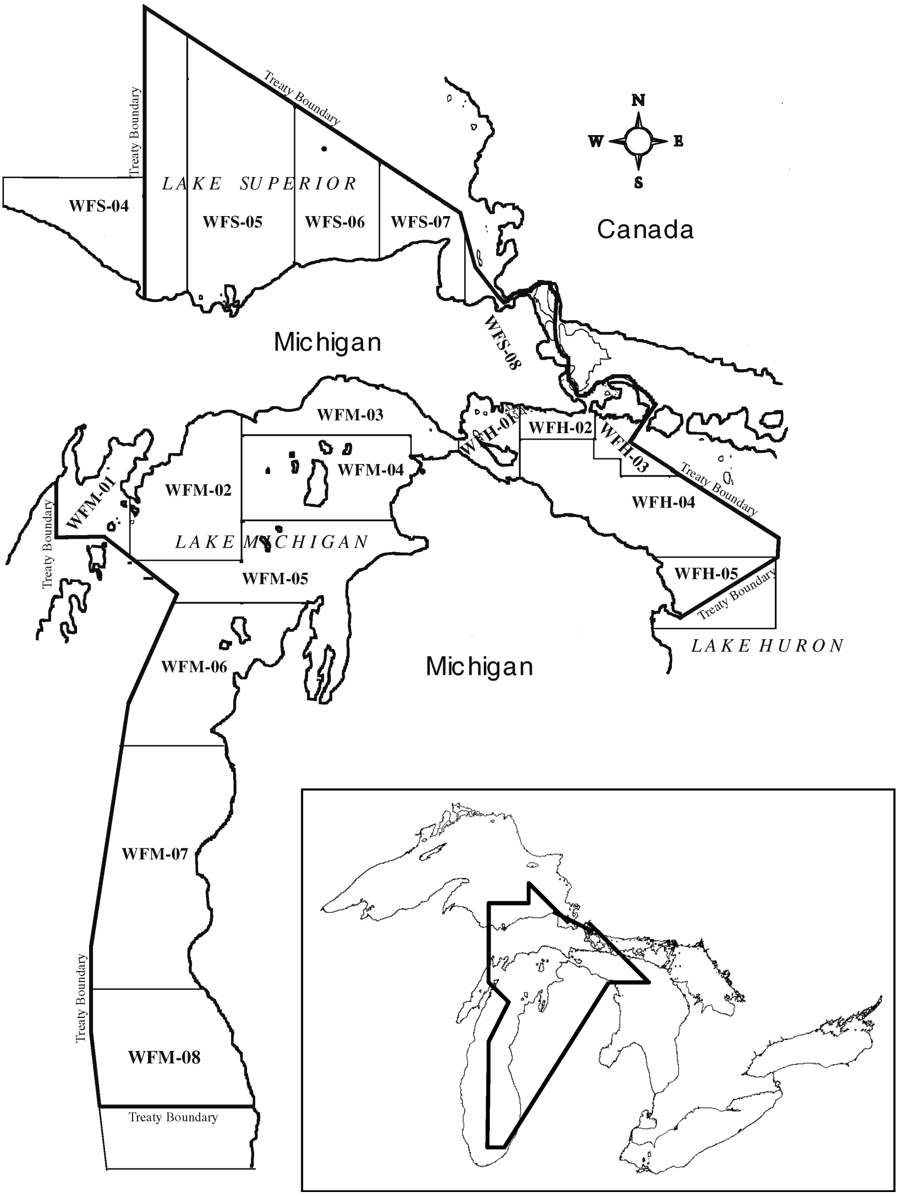

FIGURE 1.—Map of 1836 Treaty-ceded waters and lake whitefish management units in Lakes Superior (units designated as WFS), Huron (WFH), and Michigan (WFM; Ebener et al. 2005).

and specify total allowable catches (TACs) and harvest regulating guidelines (HRGs; see below). Total allowable catches are limits to catch and are used in management units where some yield is allocated to the state-licensed fishery and some to the tribal fishery. Harvest regulating guidelines are targets for yield used to guide regulations for lake whitefish in units where all yield is allocated to the tribal fishery.

The MSC fits statistical catch-at-age (CAA) models to commercial fishery data to estimate population numbers, mortality rates, fishery harvest, and other population parameters of interest. The estimates of the population parameters are used to project each stock's abundance into the future, and a TAC or HRG is then calculated by applying a reference mortality rate to the estimate of the next year's abundance.

The CAA models use fishery effort data and an assumed relationship between fishing mortality (*F*) and fishery effort. Age- and year-specific F-values are estimated as the product of age-specific selectivity (S) and year-specific fishing intensity (f) for each of two fishery gears, gill nets and trap nets:

$$F_{i,a,y} = S_{i,a} f_{i,y}, \tag{1}$$

where

$$f_{i,y} = E_{i,y} q_{i,y} \varepsilon_{i,y}, \tag{2}$$

i denotes gear type, a represents age, y denotes year, E is the fishery effort that is specific to each gear type, q represents catchability, and ε denotes multiplicative observation error. The details of the CAA models have been described by Ebener et al. (2005). Equation (2) is equivalent to assuming that the commercial fishery catch per effort (CPE; estimated as the ratio of summed aggregate catch to summed aggregate effort in each year) is, on average, proportional to average abundance over the fishing year and that deviations from this average relationship are independent variations from year to year.

Violations of the assumption that CPE is proportional to average abundance can occur if the fishing power of the gear changes or if the spatial and temporal distributions of fishery effort are nonrandom (Quinn and Deriso 1999). Violations of this assumption are called hyperdepletion (CPE declines faster than abundance at high stock sizes) and hyperstability (CPE does not decline as drastically as abundance at high stock sizes; Quinn and Deriso 1999). For example, an increase in the number of fishing operations could cause some fishermen to operate in lower-quality habitat. Thus, CPE could decline even if fish abundance did not, and hyperdepletion would therefore occur. Hyperstability is the more common occurrence and leads to overestimation of biomass and underestimation of F, which has too often gone unrecognized and led to fishery collapses (Rose and Kulka 1999; Harley et al. 2001).

To account for some of the CPE variation that is not attributable to changes in abundance and to improve assessments and associated fishery management, CPE can be standardized by fitting statistical models to the catch and effort data and then using year effect estimates as the index of abundance (Maunder and Punt 2004; Venables and Dichmont 2004). Commonly, some form of general or generalized linear model is used to standardize the CPE data (Maunder and Punt 2004). Year is usually included as one of the explanatory variables because detection of trends in abundance over time is usually the objective (Maunder and Punt 2004). Other explanatory variables often include a spatial element or some measure of individual vessel fishing power (e.g., boat size; Battaile and Quinn 2004; Bishop et al. 2004).

Our objectives were (1) to standardize lake whitefish CPE data from the upper Great Lakes to attain an index of abundance that more accurately reflects changes in lake whitefish biomass than CPE, (2) gain an improved understanding of factors that influence commercial fishery CPE for lake whitefish, and (3) compare the factors that are important for this fishery with factors that influence CPE in other fisheries of the world. For lake trout *Salvelinus namaycush* in 1836 Treaty-ceded waters, indices of abundance are currently based on the least-squares means (LSMs) for each year from a general linear mixed model (GLMM; Deroba and Bence, in press). Consequently, we explored the use of a similar GLMM for lake whitefish and compared the temporal trends in the LSMs for each year to trends in CPE. We sought to determine (1) whether LSMs account for CPE variation sources that are not considered when CPE is estimated as a ratio of summed aggregate catch to summed aggregate effort in each year and (2) whether LSMs reveal interannual trends in apparent relative abundance that are substantially different from trends based on CPE.

Methods

Study area.—Our study area encompassed the waters relevant to the 1836 Treaty, including the majority of Michigan waters of Lakes Superior, Huron, and Michigan (Figure 1). These waters were stratified into 18 management units with individual surface areas ranging from 69,000–733,000 ha and a total surface area of 5.8 million ha (Figure 1; Ebener et al. 2005). Analyses were done separately for each management unit because the lake whitefish in these units are treated as reproductively isolated stocks and because the units define the resolution of spatial stratification used to manage the species (Ebener et al. 2005).

Data and analyses.—Data were collected from commercial fishing operations as part of a requirement for all licensed vessels to submit monthly reports that describe the following information for each day of the month: weight of fish landed, amount of gear lifted, 10- × 10-min statistical grid in which the catch and effort occurred, and other auxiliary data (Ebener et al. 2005). Monofilament large-mesh gill nets (stretched mesh size $\geq$ 114 mm) and trap nets (height = 6–14 m) accounted for nearly 100% of the lake whitefish commercial harvest, and analyses were only conducted on these two gear types. The range of years included in this study differed by management unit and gear type, and data for some years were missing (i.e., no catch or effort was reported; Table 1). Analyses were only conducted on 12 of the 18 management units for the

TABLE 1.—Years of lake whitefish catch and effort data analyzed from gill-net and trap-net fisheries conducted in 1836 Treaty of Washington-ceded waters of Lakes Superior (management units designated as WFS), Huron (WFH), and Michigan (WFM).

Management unit	Gill-net years	Trap-net years
WFH-01	1981–2001	1981–1982; 1986–2001
WFH-02	1982–2001	1983; 1986–1987; 1989–2001
WFH-04	1981–2001	1981–1982; 1984–2001
WFM-01		1981–1985; 1995–1998; 2000–2001
WFM-02	1986–2001	1986–2001
WFM-03	1986–2001	1986–2001
WFM-04	1981–2001	1989–2001
WFM-05	1981–2005	1981–2001
WFM-06	1985–1989; 1993–2001	
WFS-05	1986–2001	
WFS-06	1985–2001	
WFS-07	1981–2001	1981; 1985–2001
WFS-08	1981–2002	1981–1982; 1984–1986; 1996–2001

gill-net fishery and 10 of the 18 management units for the trap-net fishery because few or no observations were recorded within most years for some management units and gears.

Catch per effort was estimated separately for gill nets and trap nets as the ratio of summed aggregate catch to summed aggregate effort in each year, as is currently used in the CAA models. Catch was measured as the round mass of lake whitefish for both gears, and effort was measured in thousands of feet of gill net or number of trap-net lifts.

General linear mixed models with $\log_e(\mathrm{CPE} + 1)$ as the dependent variable were fitted separately for gill nets and trap nets. We applied the $\log_e$ transformation because examination of the distribution of the data showed that this was necessary to meet the assumption of normality for general linear models (McCulloch and Searle 2001; Gelman and Hill 2007). We added 1.0 to all CPE observations prior to transformation to address the zero CPE observations, which were infrequent (~0.001% for each gear type). The added constant of 1.0 represents a low CPE for gill nets and the lowest possible CPE for trap nets, and more than 99% of CPE values exceeded 1.0 for both gear types.

Our initial full model for gill nets included fixed effects of year, month, and boat size, and random effects of license holder, grid, and all possible two- and three-way interactions. In preliminary analyses, higher-order interactions (four way and above) were not estimable for any management unit and so were excluded from further consideration. Because not enough individual license holders fished with multiple boat sizes, license holder and boat size were confounded when two- and three-way interactions with license holder and two- and three-way interactions with boat size were included in the same model. Furthermore, in preliminary analyses, interactions with license holder were only estimable for two management units, whereas interactions with boat size were estimable for all management units. Consequently, all interactions with license holder were also excluded from further consideration. Thus, the new full model included fixed effects of year (α_y), month (β_m), and boat size (γ_b) and random effects of license holder (c_l), grid (k_g), and all two- and three-way interactions except those with license holder:

$$\begin{aligned}\log_e(\mathrm{CPE}+1) = {} & \mu + \alpha_y + \beta_m + \gamma_b + c_l + k_g + o_{ym} \\ & + p_{yb} + q_{yg} + r_{mb} + s_{mg} + t_{bg} \\ & + u_{mbg} + d_{gmy} + h_{gyb} + j_{ymb} \\ & + \varepsilon_{iymbgl},\end{aligned} \tag{3}$$

where μ is the overall mean; o_{ym} is the interaction of year and month; p_{yb} is the interaction of year and boat size; q_{yg} is the interaction of year and grid; r_{mb} is the interaction of month and boat size; s_{mg} is the interaction of month and grid; t_{bg} is the interaction of boat size and grid; u_{mbg} is the interaction of month, boat size, and grid; d_{gmy} is the interaction of grid, month, and year; h_{gyb} is the interaction of grid, year, and boat size; j_{ymb} is the interaction of year, month, and boat size; and ε_{iymbgl} is residual error for each observation i. This model assumes that the random effects and residual error are all independent and identically distributed as normal with a mean of zero. Boat size was a categorical effect, and sizes were defined as small (≤20 ft), medium (20–30 ft), and large (≥30 ft).

The full model for trap nets included fixed effects of year and month and random effects of license holder, grid, and all two- and three-way interactions:

$$\begin{aligned}\log_e(\mathrm{CPE}+1) = {} & \mu + \alpha_y + \beta_m + c_l + k_g + o_{ym} + v_{yl} \\ & + w_{ml} + s_{mg} + x_{gl} + q_{yg} + z_{yml} \\ & + d_{gmy} + a_{ygl} + e_{mgl} + \varepsilon_{iymgl},\end{aligned} \tag{4}$$

where v_{yl} is the interaction of year and license holder; w_{ml} is the interaction of month and license holder; x_{gl} is the interaction of grid and license holder; z_{yml} is the interaction of year, month, and license holder; a_{ygl} is the interaction of year, grid, and license holder; e_{mgl} is the interaction of month, grid, and license holder; and all other terms are defined as for gill nets. In 4 of the 10 management units analyzed for the trap-net fishery, all observations came from one boat size category; therefore, the effect of boat size was not evaluated for this fishery.

Final models for both gear types were determined by evaluating which effects could be removed; this process was done using corrected Akaike's information criterion (AIC_c; Burnham and Anderson 2002). Our model selection approach was to first consider which random effects would be removed from the final model while keeping all fixed effects in the model (Ngo and Brand 1997). Random effects were selected prior to fixed effects so that the final models had the simplest error structure possible (i.e., a random effect would be eliminated rather than a fixed effect that explained similar sources of variation). Our approach to selecting random effects was to drop each random effect one at a time while keeping all other effects in the model. Once a random effect was removed, the AIC_c difference (ΔAIC_c) was then calculated by subtracting the AIC_c for the reduced model from the AIC_c for the full model. If ΔAIC_c was greater than 2.0 (Burnham and Anderson 2002), the factor that was absent from the reduced model was eliminated from the final model; otherwise, the factor was retained. We followed this approach because with 22 combinations (management unit × gear) and 12 potential random effects to consider for each combination, it was impractical to fit and compare all possible models. A random effect was also dropped from the final model if the variance estimate for that factor was zero. Restricted maximum likelihood (REML) was used for model fitting when comparing models with different random effects, given its superior performance in estimating random effects (McCulloch and Searle 2001).

Once the best set of random effects was selected, the best set of fixed effects was selected by comparing AIC_c values for all possible combinations of fixed effects. Models were fitted using maximum likelihood instead of REML, because comparisons with AIC_c based on REML are not valid for models with different fixed effects (SAS Institute 2003). During this process, the previously determined best random effects portion of the model was used. The α_y effect was not evaluated during model selection because the objective was to estimate a yearly index of abundance, and therefore it was necessary to retain year in the final model. The ΔAIC_c values are not reported in the results, because this would have required reporting a value for each factor that was included in the full models for each management unit and gear type (i.e., 298 values). Rather, we report the ΔAIC_c between a means model (i.e., a model with only a year effect) and the final model (ΔAIC_c = [AIC_c of means model] − [AIC_c of final model]) to quantify the likely improvement that each final model offers over the current abundance indices that do not account for factors other than year.

Generally, the same effects were included in the final model for each management unit, but the models for some management units could be improved by (1) the elimination of an effect that improved model fit for the majority of management units or (2) the inclusion of an effect that did not improve model fit for the majority of management units. For simplicity in reporting results of these analyses, we eliminated an effect from all management units if it only improved model fit in a minority of management units. Least-squares means for each year were calculated as the sum of μ, α_y, and the average of the coefficient estimates over all levels of fixed effects other than year in the final models (SAS Institute 2003). The LSMs for each year from the final model, as determined based on the majority of management units, were nearly identical to the LSMs from other models that improved model fit for a minority of management units. Consequently, we believe that the conclusions of these analyses are robust to this approach. However, if the estimated uncertainty (e.g., SE) associated with LSMs (or, alternatively, year effects or other functions of model parameters) is important, as is the case when fitting stock assessment models to abundance indices that are weighted (i.e., using SEs) relative to other data (e.g., Maunder 2001; Maunder and Starr 2003), a different model than the final models given here may be warranted for some management units.

Differences between back-transformed LSMs for each year and CPE + 1 were qualitatively examined by plotting the proportional difference (PD) between the two measures across years for each management unit included in this analysis. Proportional difference was calculated as

$$PD = (CPE + 1)/[\exp(LSM)]. \qquad (5)$$

The PD measures the magnitude of the difference between CPE and LSMs. For example, if PD equals 2, then the CPE is two times larger than the abundance index based on the mixed model. Since both the LSMs and CPE are relative indices, we were interested in temporal changes in PD rather than deviation of average PD from 1.0. Consequently, if PD varied without trend, we concluded that the two approaches generally suggested similar trends in abundance through time despite possible differences existing for a given year. Conversely, if PD trended through time, we concluded that the two indices suggested different temporal trends.

The relative effect of factors included in the final model on CPE was determined by averaging coefficient estimates across management units and comparing the average values. For random effects, the variance component estimates for each effect were used in estimating the average; for fixed effects, the

TABLE 2.—Difference (Δ) in the corrected Akaike's information criterion (AIC_c) between the final lake whitefish abundance model and a means model (i.e., a model with only a year effect; ΔAIC_c = means model AIC_c − final model AIC_c) for gill-net and trap-net fisheries conducted in 1836 Treaty-ceded waters of Lakes Superior (management units designated as WFS), Huron (WFH), and Michigan (WFM).

Management unit	Gill-net ΔAIC_c	Trap-net ΔAIC_c
WFH-01	536.1	281.3
WFH-02	173.9	324.8
WFH-04	508.1	121.9
WFM-01		95.1
WFM-02	93.5	28.4
WFM-03	677.7	478.2
WFM-04	556.4	212.4
WFM-05	320	2.2
WFM-06	−10.2	
WFS-05	61.1	
WFS-06	118.2	
WFS-07	993.5	111.8
WFS-08	322.1	44.4

coefficient estimates for each level of a factor were used. For boat size, the averages were estimated separately for each lake because different boat sizes may perform differently in each lake.

Results

Gill-Net Fishery

The final model for the lake whitefish gill-net fishery included fixed effects of year, month, and boat size and random effects of license holder and the year × month interaction:

$$\log_e(\text{CPE} + 1) = \mu + \alpha_y + \beta_m + \gamma_b + c_l + o_{ym} + \varepsilon_{iymbgl}. \qquad (6)$$

The final model improved model fit over a means model for all but one management unit, and the average ΔAIC_c was 362.5 (range = −10 to 2,514; Table 2). The final model did not improve fit over a means model for Lake Michigan management unit WFM-06; this management unit had the fewest number of gill-net lifts (N = 308; mean N for all units = 1,452), and therefore the data may have been insufficient to adequately capture the CPE variability caused by the various factors. Of the random effects, the license holder effect accounted for the most variation in CPE (Table 3). The effect of boat size depended on lake (Table 4). For Lake Superior, CPE did not vary much among boat size-classes. For Lake Huron, small and medium boats had similar CPEs, which were less than the CPE of large boats. For Lake Michigan, CPE was greatest for medium boats, lowest for large boats, and intermediate for small boats. Catch per effort was generally low during January–September, highest in October and November, and intermediate during December (Figure 2).

TABLE 3.—Average variance component (σ^2) estimates for a model of lake whitefish catch per effort (CPE) in the gill-net fishery (residual error = σ^2_{iymbgi}; license holder = σ^2_l; month × year = σ^2_{my}) and for a model of trap-net fishery CPE (residual error = σ^2_{iyml}; year × license holder = σ^2_{yl}; month × year = σ^2_{my}; month × year × license holder = σ^2_{myl}) for 1836 Treaty-ceded waters of Lakes Superior, Huron, and Michigan. Variance component estimates were averaged across management units.

Gear type	Variance component	Mean estimate
Gill net	σ^2_{iymbgl}	0.47
	σ^2_l	0.22
	σ^2_{my}	0.05
Trap net	σ^2_{iyml}	0.17
	σ^2_{yl}	0.29
	σ^2_{my}	0.09
	σ^2_{myl}	0.09

The index of abundance provided by the GLMMs suggested temporal patterns differing from those of CPE (i.e., PD trended through time) over some or all of the time series in some management units for the lake whitefish gill-net fishery (Figure 3). The PD for Lake Huron management units WFH-01 and WFH-04 generally varied without trend, whereas the PD in WFH-02 declined during 1982–1983 but varied without trend for the remainder of the time series. In Lake Michigan, the PD in WFM-02 increased during 1987–1988 and then decreased. In WFM-03, the PD increased in variability over the time series and increased during 1999–2001. In WFM-04, the PD generally declined through time. In WFM-05, PD generally varied without trend but declined during 1997–1999 and then increased. In WFM-06, PD declined during 1993–1997. The PD in Lake Superior management units WFS-05, WFS-06, WFS-07, and WFS-08 generally varied without trend except in WFS-05 during 1999–2001, when PD declined.

Trap-Net Fishery

The final model for the trap-net fishery included fixed effects of year and month and random effects of

TABLE 4.—Average coefficient estimates for three boat size categories (small: <20 ft; medium: 20–30 ft; large: >30 ft) in a model of lake whitefish catch per effort for the gill-net fishery in 1836 Treaty-ceded waters of Lakes Superior, Huron, and Michigan. Coefficients were averaged across management units within each lake.

Boat size	Lake Superior	Lake Huron	Lake Michigan
Large	0.03	0.11	−0.28
Medium	0.05	−0.03	0.09
Small	0.00	0.00	0.00

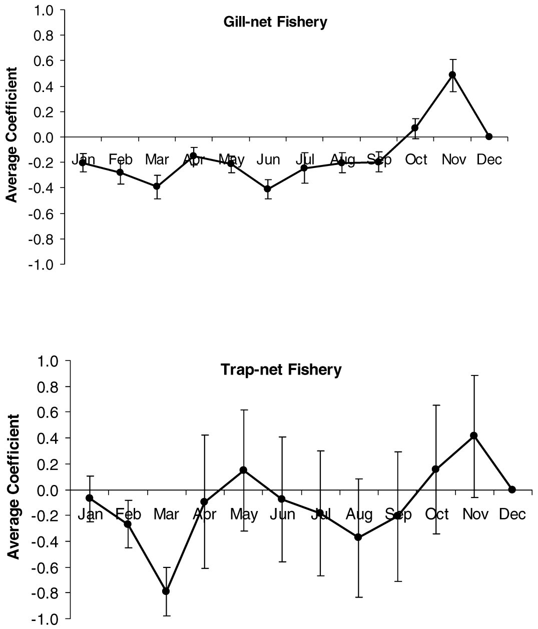

FIGURE 2.—Average coefficient estimates (±SE representing uncertainty resulting from variability among management units shown in Figure 1) for the month effect from a general linear mixed model standardizing lake whitefish catch per effort (catch = aggregate round mass of fish) for the gill-net fishery (top panel; effort = aggregate net length in thousands of feet) and trap-net fishery (bottom panel; effort = number of lifts) in 1836 Treaty-ceded waters of Lakes Superior, Huron, and Michigan. Coefficient estimates were averaged across various years (generally 1981–2001) and management units.

the month × year, year × license holder, and month × year × license holder interactions:

$$\log_e(\mathrm{CPE}+1) = \mu + \alpha_y + \beta_m + o_{my} + v_{yl} + z_{myl} + \varepsilon_{iyml}. \qquad (7)$$

The final model improved model fit over a means model in all management units by an average $\Delta\mathrm{AIC}_c$ of 170.1 (range = 2.2–478.2; Table 2). Of the random effects, the year × license holder interaction accounted for the most variation in $\log_e(\mathrm{CPE}+1)$—even more than residual error (Table 3). Catch per effort was generally lowest during January–September (with the exception of May), highest in October and November, and intermediate in December (Figure 2).

The index of abundance provided by the GLMMs suggested temporal patterns differing from those of CPE (i.e., PD trended through time) over all or some of the time series in some management units (Figure 4). For Lake Huron, the PD in WFH-01 and WFH-02 generally varied without trend; the PD in WFH-04 varied without trend until 1998, increased from 1998 to 2000, and then decreased. For Lake Michigan, the PD in WFM-01, WFM-02, and WFM-03 generally varied without trend except for a 2000–2001 increase within WFM-01. Proportional distance exhibited a 2-year cycle in WFM-04 and a 6-year cycle in WFM-05. For Lake Superior, the PD in WFS-07 generally varied without trend, whereas the PD in WFS-08 increased during 1984–1986 but varied without trend during the few other years of data.

Discussion

Catch per effort is often assumed to be proportional to abundance, but CPE can change due to factors other than abundance that cause violations of this assumption (Quinn and Deriso 1999; Battaile and Quinn 2004). Violations of the assumption of proportionality can lead to inaccurate estimates of abundance from stock assessments; in particular, hyperstability can increase the risk for fishery collapse (Rose and Kulka 1999; Harley et al. 2001). Indices of abundance based on commercial fishery catch and effort data are at an especially high risk of proportionality assumption violation due to factors such as systematic changes in characteristics of the fishing fleet (e.g., technological advancements and entrance and exit of individual vessels), nonrandom search effort, and the spatial distribution of the fish stock (Rose and Kulka 1999; Harley et al. 2001; Maynou et al. 2003; Battaile and Quinn 2004; Bishop et al. 2004; Campbell 2004). For these reasons, CPE data from many major marine fisheries are now often standardized using various statistical models (e.g., GLMMs or generalized linear models) that account for some of the CPE variation that is not attributable to abundance, thereby allowing the year effect to become a more-accurate abundance index (Maunder and Punt 2004; Venables and Dichmont 2004). Factors commonly included in models used to standardize CPE data include time (usually year), location (e.g., grid in this study), individual vessels, and characteristics of vessels that affect catchability (e.g., vessel size, horsepower, and Global Positioning System), among other factors (Maunder and Punt 2004).

The temporal trends exhibited by standardized CPE data (e.g., LSMs) have differed from those of nonstandardized CPE data (e.g., ratio of aggregate catch to aggregate effort in each year) in other studies (Maynou et al. 2003; Battaile and Quinn 2004), as was true for some management units in our evaluation of lake whitefish fisheries in the Great Lakes. Thus, we believe that model-based indices of abundance should replace nonstandardized CPE in some lake whitefish stock assessment models, especially for those management units where PD was shown to trend through time. Converting to the use of model-based indices of abundance in the stock assessment models for these management units would probably lead to more-accurate estimates (e.g., abundance) than the current approach of treating raw effort as an index of *F* (equivalent to using CPE as an abundance index). This outcome would also probably hold true for other freshwater systems, where model-based methods for standardizing CPE data have not been used as frequently as in marine systems.

Observed changes in PD in this study can be partially explained by (1) the fishing period (e.g., month) and (2) the pool of license holders from which CPE observations were drawn. For example, in the WFM-02 gill-net fishery during 1988, the spike in PD relative to other years may have been attributable to fewer observations being made during the spring (i.e., when CPE is lower relative to other times of year) and more observations being taken from license holders with relatively high CPEs. Similarly, in the WFM-04 trap-net fishery, peaks in PD occurred during years when more observations came from better-performing license holders than from other license holders. Consequently, indices of abundance based on CPE in these and other areas were probably driven by differences in the number of observations taken among seasons or by differences in license holders rather than by actual changes in abundance (i.e., an assumption of stock assessments).

In addition to providing a more-accurate index of abundance, the use of mixed-effects models also allows the uncertainty around the indices to be more accurately quantified for each year; this can be especially important when estimates of uncertainty are used to weight the importance of the yearly CPE indices in stock assessment models (Maunder and Starr 2003; Helser et al. 2004). Maunder and Starr (2003) described methods for weighting yearly indices of abundance by the coefficient of variation (CV = 100 × SD/mean) during the fitting of stock assessment models; those authors also found that stock assessment estimates (e.g., abundance) were sometimes less accurate when all yearly index values had equal weights than when year-specific weights were applied. Furthermore, Helser et al. (2004) found that ignoring the variability due to random effects, including vessel and the vessel × year interaction (similar to the effects of license holder and the license holder × year interaction in this study), could lead to an underestimation of uncertainty in indices of abundance. Thus, if the CPE data used in fitting lake whitefish stock assessment models were replaced with model-based, standardized CPE indices and associated estimates of uncertainty for each year (e.g., SEs around the LSMs), uncertainty in the indices of abundance would be more accurately quantified and CAA stock assessment estimates would also probably be more accurate. This benefit would accrue even in areas where CPE and model-based indices showed similar temporal patterns (i.e., where PD did not show any trends or systematic temporal patterns).

We do not believe that calculating a fishery CPE index by combining CPE each year over strata defined based on statistical modeling provides a viable alternative to the use of indices that are directly derived from model-based methods. This conclusion applies especially in the presence of random effects like those observed for Great Lakes lake whitefish data and that appear to be common to fishery CPE data from marine systems. A great advantage of a model-based approach is that the complex correlated error structure resulting from such random effects can be accounted for parsimoniously. The studies cited above suggest that a stratification approach would either (1) underestimate uncertainty in the indices of abundance, thus leading to inaccurate stock assessment results by ignoring variability attributable to random effects, or (2) would require so many strata with so few observations per stratum that the resulting indices would be poorly estimated. For example, our model for the gill-net fishery suggests that strata should account for seasonality, boat size, and individual license holder, but available data only consisted of monthly summaries by each license holder. Even if data were combined over similar months, few observations would be available per stratum. In some situations (e.g., if random effects are less important), data from each year could be poststratified into relatively few strata. In such a situation, calculating indices based on combining raw results over strata might be a viable approach; an advantage of this approach is that refitting of statistical models would not be necessary each time additional yearly data become available.

An alternative approach to using model-based output as an index of abundance in stock assessments is to integrate the standardization process into the estimation procedure of the stock assessment models (Maunder 2001; Maunder and Langley 2004). Such an approach models CPE data in the same manner as described here but integrates the CPE model as a submodel of the overall assessment. Maunder (2001) found that integrating the CPE standardization into the estimation procedure of the stock assessment model provided a more-accurate representation of the uncertainty in stock assessment parameter estimates. The reason for this result, however, was unclear; therefore, more research is needed in this area, especially given the programming and data management challenges associated with

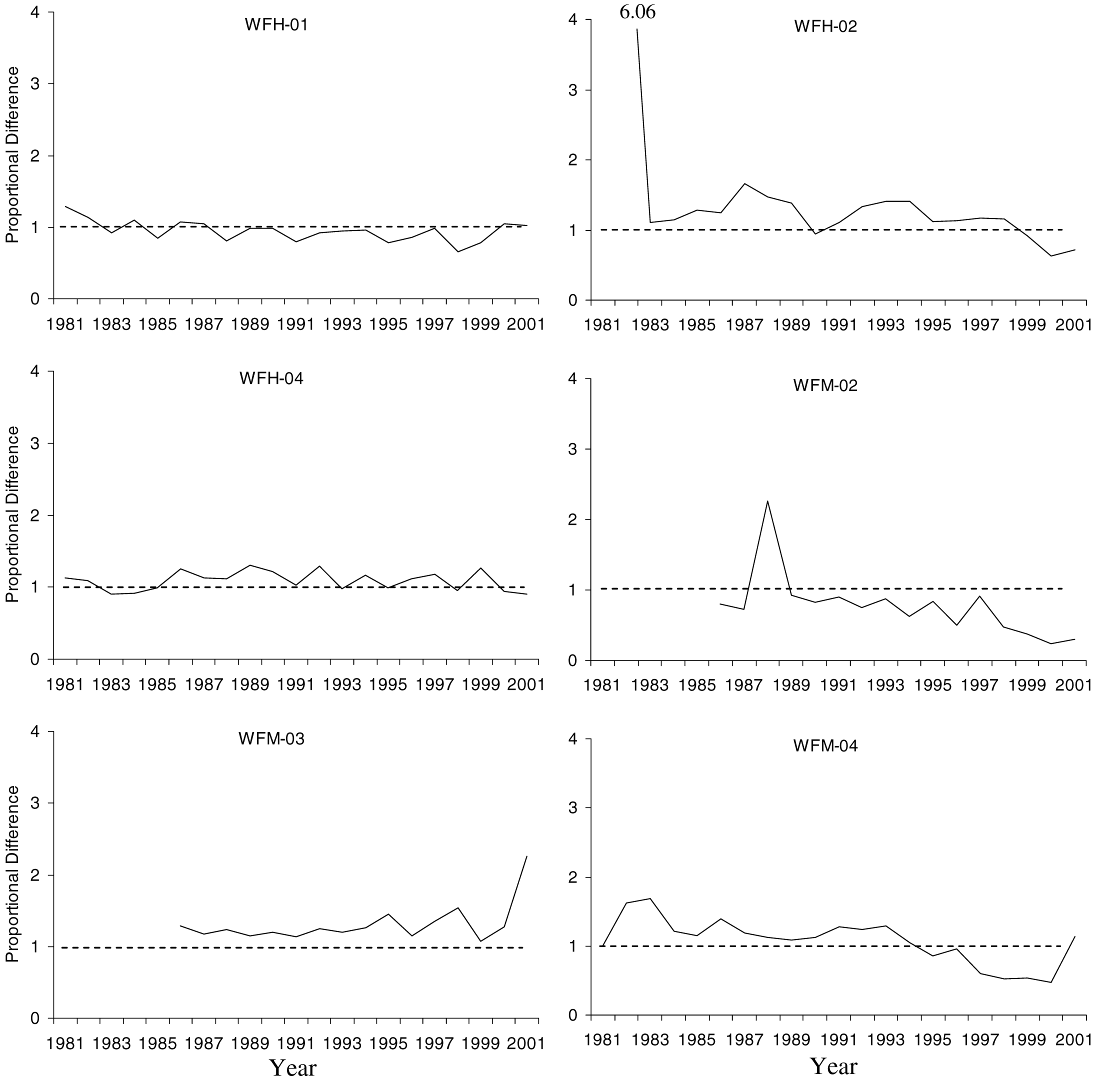

FIGURE 3.—Annual (generally 1981–2001) proportional difference between an index of lake whitefish abundance from a general linear mixed model (i.e., least-squares mean) and catch per effort (ratio of fish round mass to aggregate net length) from the gill-net fishery in 1836 Treaty-ceded waters of Lakes Superior (management units designated as WFS), Huron (WFH), and Michigan (WFM).

integrating complex GLMMs and related models for fishery CPE into assessment models.

Standardization techniques for fishery CPE data cannot ensure that every source of CPE variation other than changes in abundance has been considered. For example, changes that are confounded with year and that universally affect the fishing fleet or density-dependent changes in catchability cannot be accounted for by use of model-based standardization methods. Factors left untreated by standardization methods should be addressed in the stock assessments where the CPE abundance indices are used; one such technique is to allow for time-varying catchability (Wilberg and Bence 2006).

The factors in the final models for the gill-net and trap-net fisheries were similar to those in models developed for other fisheries (Maynou et al. 2003; Battaile and Quinn 2004; Bishop et al. 2004; Helser et al. 2004). This commonality suggests that similar factors are likely to be important and necessary for consideration when standardizing CPE data for most fisheries. Year is usually included as one of the

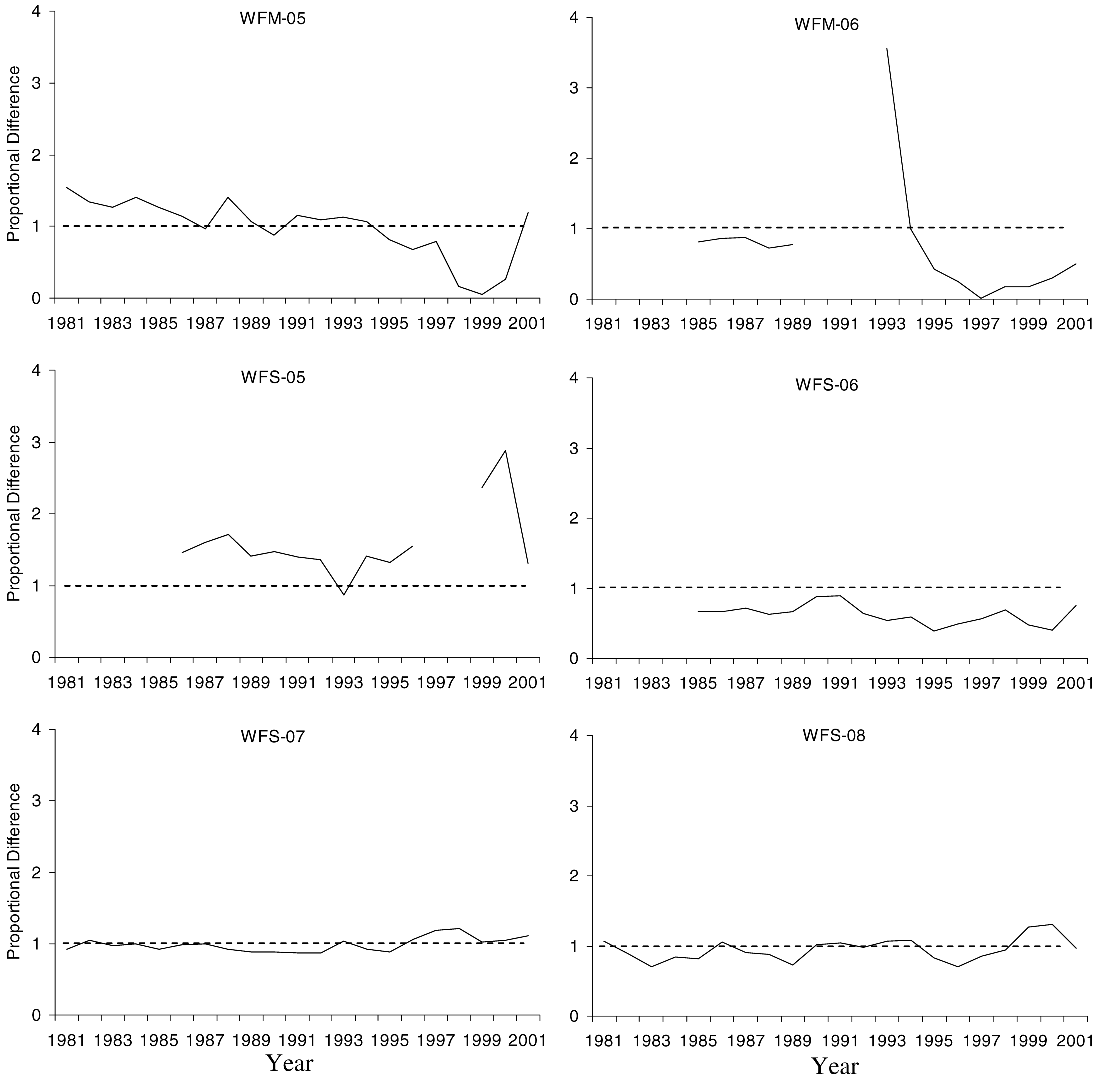

FIGURE 3.—Continued.

explanatory variables, because detecting trends through time is often the objective for developing indices of abundance (Maunder and Punt 2004), as was the case in this study. Temporal factors on a finer scale than year have also been included in statistical models used for CPE standardization to account for systematic temporal patterns in fish abundance or catchability (Battaile and Quinn 2004). Battaile and Quinn (2004) used a fixed-effects analysis of variance to standardize CPE data from the eastern Bering Sea trawl fishery for walleye pollock *Theragra chalcogramma* and found a significant effect of time of day (i.e., a categorical variable for daylight versus nighttime) wherein higher catch rates were observed during daytime than during nighttime. Battaile and Quinn (2004) suggested that the difference was attributable to greater catchability of walleye pollock during daylight, when the fish exhibit schooling behavior, than during nighttime, when the fish spread out to forage. In the present study, month was included in the final model for the gill-net and trap-net fisheries, and higher catch rates occurred from October to December, probably in association with a catchability increase facilitated by the aggregation of lake whitefish spawners in most areas of the Great Lakes during that period (Becker 1983). The results of these studies suggest that whenever possible, temporal factors accounting for systematic changes in fish

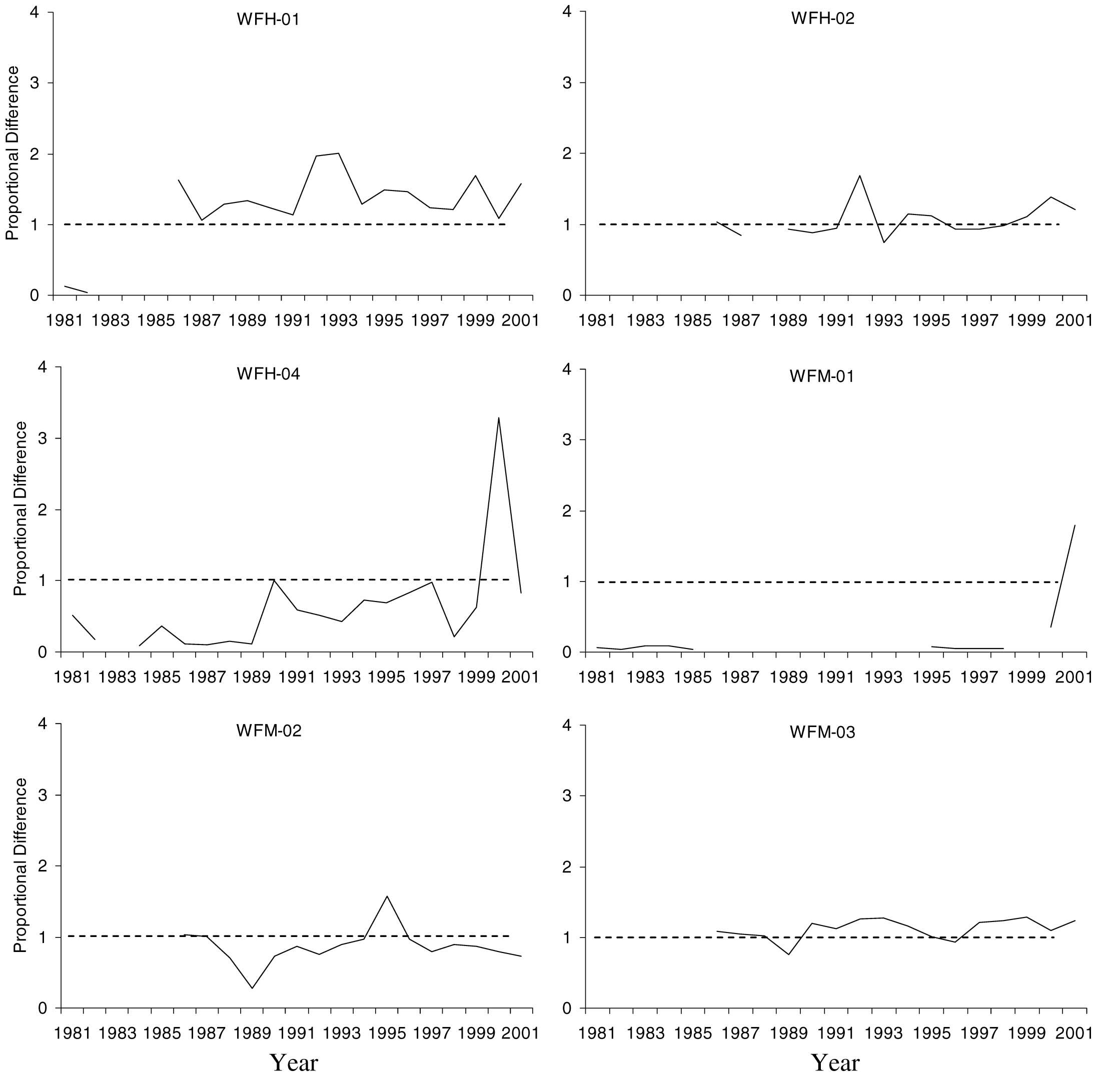

FIGURE 4.—Annual (generally 1981–2001) proportional difference between an index of lake whitefish abundance from a general linear mixed model (i.e., least-squares mean) and catch per effort (ratio of fish round mass to aggregate number of lifts) from the trap-net fishery in 1836 Treaty-ceded waters of Lakes Superior (management units designated as WFS), Huron (WFH), and Michigan (WFM).

aggregating behaviors should be considered in models used to standardize CPE data.

Various measures of vessel power have also been included in models used for standardizing CPE data. Vessel power is any measure of the boat or crew that is likely to affect catchability and the abundance indices resulting from CPE data associated with that vessel. In the walleye pollock trawl fishery of the eastern Bering Sea, longer vessels tended to have higher catch rates than shorter vessels, as indicated by the coefficient estimates for each vessel participating in the fishery (Battaile and Quinn 2004). For the trawl fishery directed at the Norway lobster *Nephrops norvegicus* and deepwater red shrimp *Aristeus antennatus* in the northwestern Mediterranean Sea, generalized linear models used for CPE standardization included measures of vessel gross tonnage, engine horsepower, and total vessel length (Maynou et al. 2003). Longer, more-powerful vessels generally had higher catch rates. In the absence of direct measures of vessel power, some surrogate measure can also be used. For example, Punt et al. (1996) included the number of crew members on the vessel as a surrogate for vessel length in generalized linear models used to standardize CPE

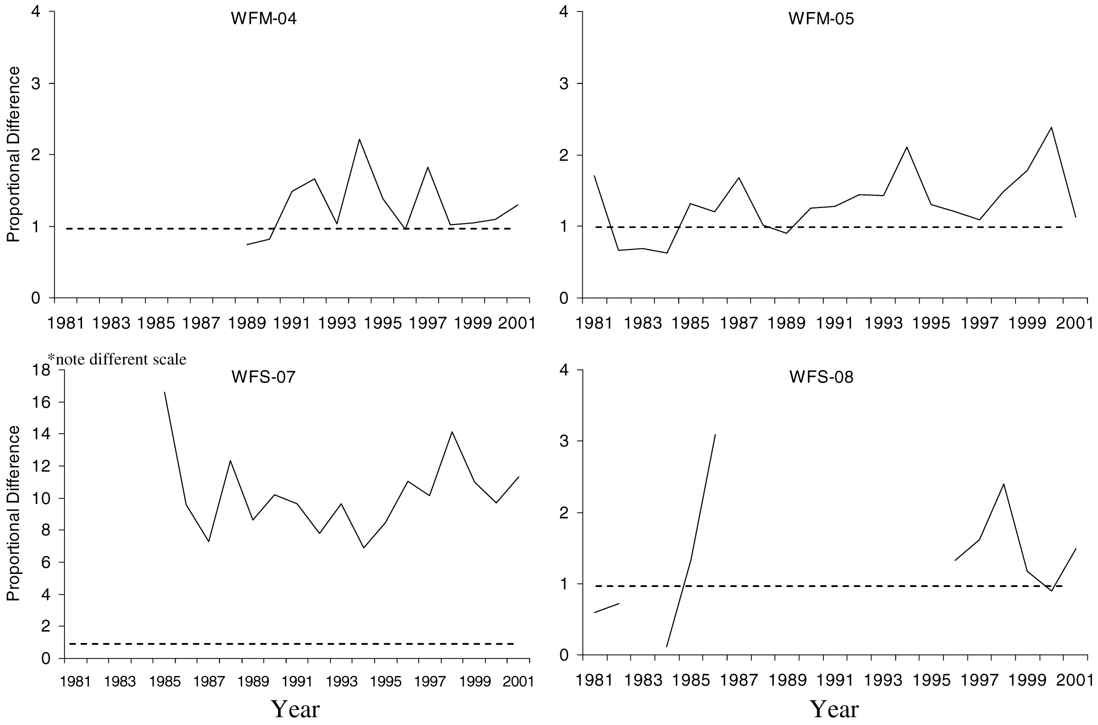

FIGURE 4.—Continued.

data from the longline fishery for albacore *Thunnus alalunga*. For the lake whitefish gill-net fishery in the present study, a categorical effect of boat size (i.e., length) was used as a measure of vessel power, but effects of boat size on CPE were inconsistent across lakes. This inconsistency hinders our ability to make broad conclusions about the relative success of various vessel sizes, but the explanation may be found in the characteristics of the lakes themselves. The depth gradient of Lake Superior is relatively steep and permits access to fishing grounds by all boat sizes, resulting in similar performance among boat sizes. Conversely, Lake Michigan offers more-shallow fishing grounds that are more accessible to small- and medium-sized boats than to large boats, possibly resulting in the relatively high catch rates for small and medium vessels. The reason for performance differences among boat sizes in Lake Huron, however, is not clear.

A factor for individual vessel, such as license holder in this study, is also commonly included in models for CPE standardization (Maynou et al. 2003; Battaile and Quinn 2004; Bishop et al. 2004; Cooper et al. 2004; Helser et al. 2004). As is reported here, an individual vessel factor explained the most variability in CPE for the eastern Bering Sea walleye pollock trawl fishery (Battaile and Quinn 2004). Generalized linear models that included vessel also explained the most variation in deepwater red shrimp CPE for the trawl fishery in the Mediterranean Sea (Maynou et al. 2003). Cooper et al. (2004) and Helser et al. (2004) found that individual vessel and interactions with vessel should be included in the final models used to standardize U.S. West Coast groundfish bottom trawl surveys. The results of Cooper et al. (2004) and Helser et al. (2004) suggest that even with survey data, the standardization of CPE is necessary and the availability of model-based indices should not replace the use of consistent survey sampling.

The consistent inclusion of an individual vessel effect indicates that this factor serves as a catch-all for boat characteristics that are not included in models (Battaile and Quinn 2004). For example, Maynou et al. (2003) suggested that the inclusion of individual vessel probably accounts for the expertise of individual fishers or unmeasured technical characteristics, such as investment in technology. The large amount of variation explained by the random effect of license holder and interactions with license holder for both fishery gears in this study also suggests that this factor accounts for the effects of some unmeasured characteristics, such as those suggested by Maynou et al. (2003).

Straightforward causal or biological mechanisms cannot be inferred for some of the two- and three-way interactions included in the final lake whitefish models. However, as Battaile and Quinn (2004) noted, identifying causal mechanisms is not required when standardizing CPE data because the purpose is to account for effects coincident with the variables included in the model. Therefore, a specific higher-order interaction does not necessarily indicate anything biologically meaningful; it may simply suggest that CPE variation coincides with variation in the combination of factors, due either to the factors themselves or other variables that covary with them.

The random effect of grid was not included in the final model for either the gill-net or trap-net fishery; this is surprising, as typically there is spatial variation in fish density or fishing success. Campbell (2004) found that nonrandomly sampled locations led to biased indices of abundance unless (1) the total habitat area of the stock was spatially stratified and (2) each CPE observation was weighted by the relative amount of sampling effort in the stratum from which the observation was taken. This result suggests that ignoring spatial variation in sampling effort can lead to biased indices of abundance. Exclusion of the grid effect from our final models could be explained by the fact that the analyses were already run on spatially stratified stocks (i.e., management units). However, the results of Campbell (2004) and the likely spatial variability in fish density and fishing success for most fisheries suggest that spatial effects must be considered when standardizing CPE data.

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