"IMPROPER SAMPLING IN NATURAL EXPERIMENTS: LIMITATIONS ON THE USE OF META-ANALYSIS RESULTS IN BAYESIAN UPDATING"

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Improper Sampling in Natural Experiments: Limitations on the Use of Meta-Analysis Results In Bayesian Updating

ABSTRACT

The natural experiment hypothesis underlying meta-analyses and their extensive designs give rise to many empty or scarcely populated cells. The implications of this improper sampling can be severe when the results are incorporated as prior knowledge in a Bayesian estimation framework. Using existing meta-analyses in marketing and a known recursive framework for updating estimates in linear regression models, the practical limitations of such priors are discussed and illustrated. Suggestions are provided to alleviate some of the problems.

Key Words: META-ANALYSIS; SAMPLING; BAYESIAN UPDATING.

1. INTRODUCTION

In recent years, researchers have come to acknowledge the need to take stock of the empirical results published over the years and see what knowledge, if any, has been accumulated. A common approach to perform this task has been meta-analysis (Glass, McGaw, and Smith 1981, Hunter, Schmidt, and Jackson 1982). In meta-analysis, independently-executed studies are viewed as imperfect replications of one overall but unplanned experiment. Analysis of variance or linear regression is then used to uncover systematic variation related to study design, data characteristics, etc. Examples in the marketing research literature are Assmus, Farley and Lehmann (1984), Farley, Lehmann, and Ryan (1982), Hite and Fraser (1988), Houston, Peter, and Sawyer (1983), Reilly and Conover (1983), Sultan, Farley, and Lehmann (1990), Churchill, et. al. (1985), Churchill and Peter (1984), and Tellis (1988).

Using the empirical study as a unit of analysis, meta-analysis has been used in two ways (Kemery, Mossholder, Dunlap 1989, Rothstein and McDaniel 1989). First, it has been used descriptively to summarise results from many studies. Second, meta-analysis has been used inferentially to make statements about relationships between variables across contexts. Correct inference, however, requires proper sampling. From the perspective of the natural experiment underlying meta-analysis, the question of whether the parameter space has been properly sampled is a relevant and important one. The consequences of a negative response to that question form the basis of the current research.

As meta-analyses rely on published results as sampling points, publication bias can give rise to improper sampling. Given self-censorship by authors and/or the publication criteria adopted by scientific journals, the reported findings within most areas of research interest are confined to a subspace of the feasible parameter space (Iyengar and Greenhouse 1988, Begg and Berlin 1988). The authors of meta-analyses generally recognise this experimental imperfectness of their designs, but seem implicitly to suggest that the problem is inconsequential to their results.

The impact of such improper sampling on meta-analysis results has seldom been addressed directly. Rust, Farley, and Lehmann (1988) try to assess the reliability and validity of inferential generalisations given publication bias as a sampling constraint. They devise a likelihood ratio method which tests for publication bias based on effect size. Applying the approach to two published meta-analyses, they find some evidence of publication bias. Given their focus on a testing methodology, however, their results fall short of assessing the impact of the identified bias.

Decisive evidence on the impact of publication bias and improper sampling in general might arise when one goes beyond the meta-analysis model itself and one starts building on the results.

One area of recent interest has been in using meta-analysis results as prior information in subsequent parameter estimation (Vanhonacker 1990). Meta-analysis generalisations updated with limited sample information on relevant time series enable stable and efficient estimation. As such, the timeliness of model-based decision making is enhanced substantially. Sultan, Farley, and Lehmann (1990), Vanhonacker (1989) and Vanhonacker, Lehmann, and Sultan (1990) contain examples where the meta-analysis results are incorporated in a formal Bayesian updating scheme. A similar direction is pursued in Montgomery and Srinivasan (1989). Such approaches require the explicit recognition of parameter heterogeneity and, hence, correct representation of that heterogeneity through proper sampling becomes paramount.

This paper addresses the issue of proper sampling in meta-analysis within an empirical Bayes updating scheme described in Vanhonacker, Lehmann, and Sultan (1990). Specifically, we assess the information content of empty or scarcely populated cells available for enhancing parameter stability and efficiency. In doing so, we establish limitations on the practical usefulness of meta-analysis results as prior information in light of sparcity of empirical results.

2. BASIC METHODOLOGY

1. Review and Discussion

In a meta-analysis, parameter estimates of past studies are the dependent observations. The literature is searched for empirical values on a parameter of interest. Tellis (1988), for example, accumulated 421 values for price-response parameters. A set of variables are defined subsequently which might explain systematic differences in these estimates across studies. Those independent variables are viewed as the design variables of the natural experiment. In other words, parameter estimates from the past studies are viewed as sample points within the "imperfect" experiment whose design is defined by the independent measures. These independent variables can contain design characteristics of the original studies (such as, e.g., time period of observation, estimation method used, model specification), data characteristics (such as, e.g., data interval, sample size, national setting), and other variables. Interpreted alternatively, these variables specify design cells and identify to which cell each of the past estimates belongs.

Whether the analysis is performed with ANOVA (as in, e.g., Assmus, Farley, and Lehmann 1984) or regression (as in, e.g., Tellis 1988), a basic linear model underlies the methodology. The model can be expressed as (using matrix algebra notation)

$$\hat{\beta} = Zc + \varepsilon \tag{1}$$

where $\hat{\beta}$ is a column vector which contains the parameter estimates of the previous studies, Z is a design matrix, c is the parameter vector pertaining to the design variables, and ε is a vector of disturbances. The columns of matrix Z correspond to the design variables theoretically motivated to capture systematic variation in the published estimates. Accordingly, each row contains the design profile of the past study whose parameter estimate is the corresponding observation in vector $\hat{\beta}$. That row vector also identifies the cell of the natural experiment to which the corresponding study belongs. Hence, grouping of the studies on profile similarity gives a quick reading of sampling frequencies. Note that in the tradition of experimental design, the columns of Z are orthogonal to one another.

The disturbance term ε has a complex structure as it recognizes two error components: one is the error estimating the dependent observations (i.e., $\hat{\beta}_i - \beta_i$ for all i); the other is the sampling error of the individual studies relative to the cell means (i.e., $\beta_i - \bar{\beta}_j$ for all i and j) where j denotes the cell to which past study i belongs). The latter component characterises the random coefficient assumption implicit in meta-analysis. Note that the systematic variance captured by Z is the systematic variance in those cell means ; in other words, the cell means are a linear combination of the columns in Z, or

$$\bar{\beta} = Zc$$

where $\bar{\beta}$ is a column vector which contains for each original study the corresponding cell mean within the natural design (i.e., for all studies belonging to the same cell - having identical corresponding rows in Z - the corresponding entries in $\bar{\beta}$ are identical).

The statistical properties of ε are such that $E(\varepsilon) = 0$ and $E(\varepsilon\,\varepsilon') = \Delta$ where E denotes the expectation operator. When the meta-analysis focuses on a single parameter (as is commonly the case in marketing), Δ is a diagonal variance-covariance matrix with as diagonal elements the sum of the two error components discussed above. When the meta-analysis is done simultaneously on a number of parameters, then matrix Δ is block diagonal with each block being the sum of the variance-covariance matrices corresponding to the two error components (Vanhonacker, Lehmann, and Sultan 1990). With these disturbance characteristics, Generalised Least Squares (GLS)

estimates can be obtained for c as $\hat{c} = (Z'\Delta^{-1}Z)^{-1} Z'\Delta^{-1} \hat{\beta}$. These estimates measure the extent to which the selected design variables contained in the columns of Z describe the systematic variance in the parameter estimates of the past studies.

This GLS estimate, $\hat{c}$, can be used as prior information on the parameter of interest in a new study. Suppose y_0 and X_0 contain, respectively, the sales and sales predictor observations in a new market response model. In matrix algebra notation the response model which is linear in parameters, can be expressed as $y_0 = X_0\beta_0+u_0$. One column in X_0 might, for example, contain advertising observations. Partitioning the matrix X_0 to separate out that particular column, we obtain $X_0=[X_{01} \vdots x_{02}]$ where column vector x_{02} contains the advertising observations. Partitioning the parameter vector β_0 accordingly, we have $\beta_0' = [\beta_{01}' \vdots \beta_{02}]$ where β_{02} denotes the true advertising response parameter in the new market response model. Accordingly, the response model can be expressed as

$$y_0=X_{01}\beta_{01}+x_{02}\beta_{02}+u_0.$$

If a meta-analysis previously had been done on the advertising response parameter (e.g., the advertising elasticity in Assmus, Farley, and Lehmann 1984), then $\hat{c}$ of the meta-analysis could be used to provide a prior value for the advertising response parameter in the new sales response study. Specifically, the prior $b_{02}= z_0\hat{c}$ could be obtained where z_0 is a row vector which describes the profile of the new study in terms of the design variables contained in Z.[1] This prior can be written as $b_{02}= z_0D\hat{\beta}$ where $D = (Z'\Delta^{-1}Z)^{-1} Z'\Delta^{-1}$, which can be expressed further as

$$b_{02}= z_0 De \beta_{02}+ z_0 D \beta_d+ u \tag{2}$$

where e denotes a vector containing all ones, β_{02} denotes the true advertising parameter value, β_d denotes a vector containing the difference between the mean advertising parameter value for the cell corresponding to a specific past study and the mean of the cell to which the new study belongs (hence, for past studies whose profile is identical to z_0, the corresponding entries in β_d are zeros), and u denotes a disturbance term. It is evident from expression (2) that $(b_{02}-z_0D\beta_d) = z_0De\beta_{02}+u$ which for $b_{02}^* = (b_{02}-z_0D\beta_d)$ can be incorporated in an expanded system as

$$\begin{bmatrix} y_0 \\ b_{02}^* \end{bmatrix} = \begin{bmatrix} X_{01} \\ 0 \end{bmatrix} \beta_{01} + \begin{bmatrix} x_{02} \\ z_0De \end{bmatrix} \beta_{02} + \begin{bmatrix} u_0 \\ u \end{bmatrix}. \tag{3}$$

An updated estimate of β_{02} can be obtained using GLS on (3).

The estimate of the entire parameter vector incorporating the prior information b_{02}^* on parameter β_{02} (i.e., GLS on (3)) can be expressed as

$$\begin{bmatrix} \hat{\beta}_{01} \\ \hat{\beta}_{02} \end{bmatrix} = \begin{bmatrix} X_{01}'\Sigma_0^{-1}X_{01} & X_{01}'\Sigma_0^{-1}X_{01} \\ x_{02}'\Sigma_0^{-1}X_{01} & x_{02}'\Sigma_0^{-1}x_{02} + \left[(z_0De)^2/\sigma_u^2\right] \end{bmatrix}^{-1} \begin{bmatrix} X_{01}'\Sigma_0^{-1}y_0 \\ x_{02}'\Sigma_0^{-1}y_0 + \left[(z_0De)^2/\sigma_u^2\right]b_{02}^* \end{bmatrix} \tag{4}$$

where $E(u_0u_0') = \Sigma_0$ and $E(u^2) = \sigma_u^2$. The simple sample GLS estimate without prior information on parameter β_{02} equals

$$\begin{bmatrix} \tilde{\beta}_{01} \\ \tilde{\beta}_{02} \end{bmatrix} = \begin{bmatrix} X_{01}'\Sigma_0^{-1}X_{01} & X_{01}'\Sigma_0^{-1}x_{02} \\ x_{02}'\Sigma_0^{-1}X_{01} & x_{02}'\Sigma_0^{-1}x_{02} \end{bmatrix}^{-1} \begin{bmatrix} X_{01}'\Sigma_0^{-1}y_0 \\ x_{02}'\Sigma_0^{-1}y_0 \end{bmatrix}. \tag{5}$$

Combining expressions (4) and (5), it can be shown that

$$\begin{bmatrix} \hat{\beta}_{01} \\ \hat{\beta}_{02} \end{bmatrix} = \begin{bmatrix} \tilde{\beta}_{01} \\ \tilde{\beta}_{02} \end{bmatrix} + \begin{bmatrix} X_{01}'\Sigma_0^{-1}X_{01} & X_{01}'\Sigma_0^{-1}x_{02} \\ x_{02}'\Sigma_0^{-1}X_{01} & x_{02}'\Sigma_0^{-1}x_{02} \end{bmatrix}^{-1} \begin{bmatrix} 0 \\ \left[(z_0De)^2/\sigma_u^2\right]\left(b_{02}^* - \hat{\beta}_{02}\right) \end{bmatrix}$$

from which it can be shown that the posterior (updated) estimate $\hat{\beta}_{02}$ in (4) can be written compactly as

$$\hat{\beta}_{02} = \left[1 + \delta\left[(z_0De)^2/\sigma_u^2\right]\right]^{-1}\left[\tilde{\beta}_{02} + \delta\left[(z_0De)^2/\sigma_u^2\right]b_{02}^*\right] \tag{6}$$

where the scalar δ equals

$$\delta = \left(x_{02}'\Sigma_0^{-1}x_{02}\right)^{-1} + \left(x_{02}'\Sigma_0^{-1}x_{02}\right)^{-1}x_{02}'\Sigma_0^{-1}X_{01}\left[X_{01}'\Sigma_0^{-1}X_{01} - X_{01}'\Sigma_0^{-1}x_{02}\left(x_{02}'\Sigma_0^{-1}x_{02}\right)^{-1} x_{02}'\Sigma_0^{-1}X_{01}\right]^{-1}X_{01}'\Sigma_0^{-1}x_{02}\left(x_{02}'\Sigma_0^{-1}x_{02}\right)^{-1}.$$

Using a matrix-inversion identity suggested in Bartlett (1951), it can be shown that δ equals

$$\delta = 1 \Big/ \left[x_{02}' \left[\Sigma_0^{-1} - \Sigma_0^{-1} X_{01} \left(X_{01}' \Sigma_0^{-1} X_{01} \right)^{-1} X_{01}' \Sigma_0^{-1} \right] x_{02} \right].$$

Plugging this expression into (6), it is evident that the posterior estimate $\hat{\beta}_{02}$ equals a weighted average of the simple sample estimate $\tilde{\beta}_{02}$ and the prior b_{02}^*. The weight given to the prior is proportional to the variance of the sample estimate (which equals $1/\delta$) over the variance of the prior. Hence, consistent with intuition and previous results (see, e.g., Leamer 1978), the prior gets more weight as it is estimated more efficiently relative to the simple sample estimate. The proportionality factor $(z_0De)^2$ will be discussed shortly.

2. What information is contained in the prior value b_{02}^* ?

The prior value b_{02}^* incorporated in (3) is based on the expression

$$b_{02}^* = z_0 \, D \, e \, \beta_{02} + u$$

and, given expression (2)

$$b_{02}^* = b_{02} - z_0 \, D\beta_d. \tag{7}$$

Accordingly, b_{02}^* is an adjustment over b_{02}. Since $D\beta_d$ can be interpreted as a least squares estimate of the design matrix Z on β_d, the predicted value b_{02} is adjusted for differences in cell means. As will become evident momentarily, the adjustment is in a direction which is of particular interest given the research question addressed here.

By definition

$$\beta_d = \bar{\beta} - \bar{\beta}_j e$$

where $\bar{\beta}$ is a column vector containing the means of the cells to which each corresponding past study belongs, $\bar{\beta}_j$ is the mean for the cell to which the new study belongs, and e is a vector containing all ones.

Since $\bar{\beta} = Zc$ is implied by a meta-analysis design, we can express b_{02}^* as

$$b^*_{02} = b_{02} - z_0 c + \bar{\beta}_j z_0 De.$$

Now, since $E(b_{02}) = z_0 c$ when $\hat{c}$ is an unbiased estimator of c, we have

$$E\left(b^*_{02}\right) = \bar{\beta}_j z_0 De . \tag{8}$$

When the design matrix Z contains a constant (or intercept) in the first column, e can be written as

$$e = Z \begin{bmatrix} 1 \\ 0 \\ \vdots \\ 0 \end{bmatrix} .$$

Accordingly, since

$$z_0 De = z_0 \left(Z' \Delta^{-1} Z\right)^{-1} Z' \Delta^{-1} e$$

we have

$$z_0 De = z_0 \begin{bmatrix} 1 \\ 0 \\ 0 \\ \vdots \\ 0 \end{bmatrix}$$

which implies that $z_0 De = 1$.

Incorporating this into (8), it is evident that

$$E\left(b^*_{02}\right) = \bar{\beta}_j .$$

In words, when Z contains a constant, or intercept (as is commonly the case in market response models), the expected value of b^*_{02} is equal to the mean parameter value of the cell to which the new study belongs. Moreover, since $z_0 De$ equals 1 in that instance, the posterior estimate $\hat{\beta}_{02}$ in (6) equals

$$\hat{\beta}_{02} = \left[1 + \left[\mathrm{Var}\left(\tilde{\beta}_{02}\right) / \sigma^2_u\right]\right]^{-1} \left[\tilde{\beta}_{02} + \left[\mathrm{Var}\left(\tilde{\beta}_{02}\right) / \sigma^2_u\right] b^*_{02}\right] \tag{9}$$

and the expanded system in (3) is essentially a straightforward Goldberger-Theil estimator (see, e.g. Rao and Yamada 1988) <u>where the prior is the mean of the cell to which the new study belongs</u>. This result is intuitive as the relevant information is confined to that cell. Nevertheless, this has not been fully recognised in previous studies.

This result raises serious questions in relation to the scarcity issue charaterizing unplanned natural experiments. For example, what does one do when the new study belongs to an empty cell (and, hence, the mean cannot be estimated)? Furthermore, what happens when improper sampling gives rise to small sample biases in the estimate $\hat{c}$? These questions are addressed in some detail hereafter.

3. IMPROPER SAMPLING

One of the characteristics of an unplanned natural experiment such as the one underlying a meta-analysis is the number of empty and scarcely populated cells. Table 1, for example, illustrates this for the meta-analyses reported in Assmus, Farley, and Lehmann (1984) and Tellis (1988). As is evident from the numbers, the large majority of cells in the design are empty; hence, there is no sample information for extensive regions of the parameter space. Furthermore, those cells that do contain observations are sampled unevenly. For example, in Tellis (1988) 52 out of 421 observations (or 12% of the total sample) belong to the same cell. This leaves 369 observations for 208 remaining cells. Obviously, a large number of cells are scarcely populated; hence, there is some likelihood of improper sampling within these cells.

In order to demonstrate the consequences of this situation, it is important to point out that the GLS estimate $\hat{c}$ can be expressed as a matrix-weighted average of subsample estimates. Specifically, for an arbitrary split of observations such as $Z' = [Z_1' \; Z_2' \; Z_3']$ and $\hat{\beta}' = [\hat{\beta}_1' \; \hat{\beta}_2' \; \hat{\beta}_3']$, knowing that

$$E(\varepsilon\varepsilon') = \Delta = \begin{bmatrix} \Delta_1 & 0 & 0 \\ 0 & \Delta_2 & 0 \\ 0 & 0 & \Delta_3 \end{bmatrix},$$

the GLS estimate $\hat{c}$ can be expressed as

$$\hat{c} = \left[Z_1' \Delta_1^{-1} Z_1 + Z_2' \Delta_2^{-1} Z_2 + Z_3' \Delta_3^{-1} Z_3\right]^{-1}$$

$$\left[Z_1' \Delta_1^{-1} Z_1 \; \hat{c}_1 + Z_2' \Delta_2^{-1} Z_2 \; \hat{c}_2 + Z_3' \Delta_3^{-1} Z_3 \hat{c}_3\right] \qquad (10)$$

where

$\hat{c}_i = \left(Z_i' \Delta_i^{-1} Z_i\right)^{-1} Z_i' \Delta_i^{-1} \widehat{\beta}_i$ for i = 1,2,3. Note that since the variance-covariance matrix of GLS estimate $\hat{c}_i$ (for i = 1,2,3) equals $\left(Z_i' \Delta_i^{-1} Z_i\right)^{-1}$, the weights in equation (10) are proportional to the inverse of the corresponding variance-covariance matrix.

Empty and scarcely populated cells are likely to give rise to small sample biases despite the fact that theoretically the $\hat{c}$ estimate is unbiased. To see that the latter is true, consider the case that $\overline{Z} = \left[Z_1' \; Z_2'\right]$ and $\hat{\overline{\beta}}' = \left(\widehat{\beta}_1' \; \widehat{\beta}_2'\right)$. In relation to the partitioning from which (10) is derived, both Z_3 and $\widehat{\beta}_3$ are unobserved (i.e., corresponding to empty cells).

The corresponding GLS estimate for c equals

$$\hat{c}_b = \left(\overline{Z}' \overline{\Delta}^{-1} \overline{Z}\right)^{-1} \overline{Z}' \overline{\Delta}^{-1} \hat{\overline{\beta}} \tag{11}$$

where $\overline{\Delta}$ is the corresponding submatrix of $E\left(\varepsilon \varepsilon'\right)$. Note that $[I \;\; 0] \widehat{\beta} = \hat{\overline{\beta}}$ and, hence,

$$\hat{c}_b = \left(\overline{Z}' \overline{\Delta}^{-1} \overline{Z}\right)^{-1} \overline{Z}' \overline{\Delta}^{-1} [I \;\; 0] \widehat{\beta} .$$

Substituting $\widehat{\beta}$ with expression (1) and taking expectations, it is evident that $E\left(\hat{c}_b\right) = c$ since $E(\varepsilon) = 0$ (and, hence, $E([I \;\; 0]\, \varepsilon) = 0$). Thus, even when some cells are empty the GLS estimate of c is unbiased. This does not imply, however, that $\hat{c}_b$ in (11) equals $\hat{c}$ in (10). It is easily demonstrated that

$$\hat{c} = \left[\sum_{i=1}^{3} \left(Z_i' \Delta_i^{-1} Z_i\right)\right]^{-1} \left[\sum_{i=1}^{2} \left(Z_i' \Delta_i^{-1} Z_i\right) \hat{c}_b + Z_3' \Delta_3^{-1} Z_3 \hat{c}_3\right].$$

That is, the GLS estimate based on full and proper sampling of the design space equals a matrix-weighted average of the GLS estimate based on the limited sample $(\hat{c}_b)$ and the GLS estimate which could be derived from observations out of the unsampled region of the design space $(\hat{c}_3)$. Whenever the latter is non-zero, it follows that $\hat{c} \neq \hat{c}_b$.

Furthermore, the fact that both $\hat{c}$ and $\hat{c}_b$ are unbiased does not imply that either equals c for any particular sample. It is reasonable to assume, however, that because of full and proper sampling, $\hat{c}$ is highly efficient which would make it a more accurate estimate of c than is $\hat{c}_b$. Moreover, improper sampling could give rise to small sample biases and more so as $\hat{c}$ becomes a more efficient estimate of c than $\hat{c}_b$.[2] The latter condition is merely academic as full and proper sampling could be driven to the limit so that $\hat{c}$'s efficiency is guaranteed. In sum, we have

established that $\hat{c}$ and $\hat{c}_b$ are (a) unbiased, and (b) generally unequal, and we make the argument that $\hat{c}_b$ is less efficient and is more likely to exhibit small sample biases than is $\hat{c}$.

Small sample biases will affect the posterior estimate $\hat{\beta}_{02}$ in (6) in two ways: (1) through the prior value directly, and (2) through the efficiency (and, hence, the relative weight) of the prior value. The prior value itself will be affected because sampling bias is carried over. Note that when the new study belongs to an empty cell (i.e., z_0 is different from any row in Z), b^*_{02} in (7) cannot even be computed. In this instance, alternative ways to derive a prior value will have to be pursued. One approach, advocated in Sultan, Farley, and Lehmann (1990), is to use the predicted value $b_{02} = z_0 \hat{c}_b$ which is the prior value assuming complete exchangeability (Vanhonacker, Lehmann, and Sultan 1990). An updated estimate for β_{02} can then be derived incorporating that value into expanded model (3) as

$$\begin{bmatrix} y_0 \\ b_{02} \end{bmatrix} = \begin{bmatrix} X_{01} \\ 0 \end{bmatrix} \beta_{01} + \begin{bmatrix} x_{02} \\ 1 \end{bmatrix} \beta_{02} + \begin{bmatrix} u_0 \\ v \end{bmatrix} \tag{12}$$

where v is a disturbance term. Using GLS on (12) an updated estimate for β_{02} can then be derived. As this posterior estimate is a matrix-weighted average of the sample estimate and b_{02} (as in equation (9)), the sample bias in $\hat{c}_b$ will affect the GLS estimate of parameter β_{02}.

Second, the weight applied to the prior value in deriving the posterior estimate will be affected by small sample biases. As $\hat{c}$ is likely to be a more efficient estimate than $\hat{c}_b$ (i.e., $\left[\left(\bar{Z}' \bar{\Delta}^{-1} \bar{Z}\right)^{-1} - \left(Z' \Delta^{-1} Z\right)\right]$ is a positive definite matrix), the prior value will be given less weight relative to the independent sample estimate $\tilde{\beta}_{02}$ as shown in expression (5). Fortunately, it is a weight reduction which will limit to some extent the impact of the biased prior. It will not, however, eliminate the bias and erroneous insights and/or conclusions that could be drawn from the posterior estimate.

4. DESIGN REDUCTION

Improper sampling in a natural experimental design is difficult to address directly as the researcher has no control over the sampling. The question arises whether something could be done to guard oneself against small sample biases. One approach might be design reduction as this affects the distribution of sampling points across the (fewer) design cells. Through reducing the natural experimental design, one could (a) ensure that all cells in the final design have

observations, and (b) potentially enhance proper sampling within each of the retained cells. This approach is discussed in some detail hereafter.

There are two ways to reduce the design of a meta-analysis. One is to recode the design variables in such a way as to collapse some levels while retaining all design variables. The second way is to delete some of the design variables entirely (i.e., reduce the number of columns in Z). This latter approach is of particular interest here because (a) it does not alter the fundamental nature of the design (i.e., one looks at the projection of the original design into a subspace), and (b) deleting design variables will not bias the parameter estimates of the remaining design variables because these variables are created to be orthogonal.

In deciding on which design variables to drop and when to stop, one has to keep the reduction objective in mind. The problem with the full design is that prior parameter estimates might exhibit small sample biases because of either no sampling points in the cell corresponding to the new study or improper sampling of the entire experimental space. To the extent that the prior estimate receives weight in the posterior estimate the biases would be carried through, possibly resulting in erroneous insights about the new study. The objective of design reduction is to limit such unfavorable results. There is no guarantee that biases will be prevented altogether, since deletion of design variables can correct the problem of zero or few observations in relevant cells but will not necessarily correct for improper sampling within any given cell.

Given this constraint, the question arises how the elimination process should be performed. It seems intuitively logical to sequentially drop design variables in terms of their parameter significance in the meta-analysis model. Note, however, that we start from a recovered design and, given a prespecified level of Type I error (α - level), all design variables will be statistically significant. Nevertheless, among the variables there will be varying degrees of significance, and the least statistically significant ones should be dropped first. Dropping these variables one at a time will reduce the number of cells in the natural design quite rapidly and potentially will enhance sampling in the remaining cells. The sample estimates of the parameters of the remaining design variables will change slightly (although they remain unbiased as indicated above), as the less efficient variables are dropped first. As the relative efficiency of the remaining variables increases, the sample estimates will become more stable and anchor around a specific value (see Appendix 1).

As for a stopping rule, we propose that reduction should continue until sampling within the cell to which the new study belongs is judged adequate. The more sampling points within the corresponding cell the more likely proper sampling of relevant information has been accomplished. Hence, the stopping point is a judgment call on the part of the researcher and will be different

depending on which new situation is being faced (i.e., new studies are likely to have different profiles within the reduced designs and, hence, will rely on different prior information). It should again be stressed, however, that increasing the number of sampling observations does not necessarily imply that the resulting sample will be proper; hence,the proposed reduction approach has the potential for enhancing results but does not guarantee it. An empirical example of design reduction and its implications on posterior estimation is discussed next.

5. EMPIRICAL ILLUSTRATION

The original data bases underlying the meta-analysis on advertising elasticities by Assmus, Farley, and Lehmann (1984) and the meta-analysis on price elasticities by Tellis (1988) were obtained to illustrate the approach in a rich empirical framework. The meta-analysis reported in Assmus, Farley, and Lehmann (1984) incorporated 128 advertising elasticity estimates and 25 design variables. Using their original coding scheme, the regression results are shown in Table 2. That table also contains the results for a reduced design (where blank spaces indicate that a variable was dropped) and the profile vector (z_0) of a new data set under the original and reduced designs. The reduced design was obtained by deleting the least statistically significant variables in the original (full) design. The stopping rule was such that the cell to which the new data set belonged had one prior observation. In other words, the reduced design was such that design matrix Z had at least one row identical to z_0. The meta-analysis reported in Tellis (1988) incorporated 422 price parameter estimates and 21 design variables. The regression results using the original coding scheme are contained in Table 3.[3] Again, a reduced design was estimated using the approach discussed above. Overall, the results are in line with what was reported in the original studies. There are some minor deviations which are likely to be a direct result of the coding scheme and/or the sample used. For example, Tellis (1988) confined himself to 368 "usable" results whereas the estimates in Table 3 are based on 421 prior estimates (i.e., one prior study of the entire data base was dropped because the price response parameter was missing). For the substantive interpretation of the results in Tables 2 and 3, the reader can refer to the original publications.

The prior values derived for the price and advertising elasticities are shown in Table 4. The first set are the predicted priors $z_0\hat{c}$, using the profile z_0 within the full design for the new data set and the $\hat{c}$ estimate. The advertising elasticity value of -0.3051 and the price elasticity of 1.8835 have no face validity. For the reduced design, the b^*_{02} values of 0.0325 for the advertising elasticity and -2.3360 for the price elasticity have the correct sign and seem intuitively reasonable. These priors play a critical role in the recursive estimation shown in (6). They are the anchoring points from which sample information will be added to adjust the estimates in a particular direction. Hence, irrespective of whether they are biased or not they will impact the recursive estimate

particularly when sample information is limited. The recursive estimator does not improve or correct on the statistical properties of the prior; the latter is simply considered a starting point. Given the dubious value of some prior estimates (particularly those predicted from the meta-analysis regression in Table 4), misleading estimates will be derived until the sample information has totally discounted the prior. If the variability in the sample information is substantial this might never happen. This case is amply illustrated below. Thus, considering only the prior values, the empirical evidence already is in favor of design reduction and recursive estimation according to (3) (i.e., confining prior information to the cell to which the new study belongs).

Table 5 shows recursive estimates as well as Ordinary Least Squares (OLS) estimates obtained by augmenting the sample for the new case, one observation at a time. No prior information is incorporated in the OLS estimates; they are shown here only as bench marks for the updated estimates. The objective of the recursive estimation methodology is to enhance parameter stability and convergence of the estimates towards the long-run value represented by the OLS sample estimates.

For price-elasticity, the updated estimates remain locked around the prior value irrespective of the number of sample observations incorporated. As the inefficient OLS estimates indicate (i.e., large estimated variances), the sample observations contain little information on that parameter. In other words, price is an insignificant predictor and, hence, in Bayesian updating the prior is carried forward. Although technically correct this result generates concern, particularly since the recursive price elasticity estimates are found to be statistically significant. This highlights a basic problem of Bayesian estimation given incorrect priors, in which instance the expected squared error loss can be substantial (Leamer 1978). However, the alternative is that no estimate is obtained altogether.

The updated advertising elasticity estimates are immediately close to the long-run estimate and remain quite stable given varying sample sizes. With few sample observations, the updated estimates are substantially smaller than the corresponding sample estimates using no prior information. These results are graphically summarised in Figure 1. That figure also illustrates the estimates derived from (11) for $b_0 = z_0 \hat{c}$ with $\hat{c}$ derived from the original full design in Assmus, Farley, and Lehmann (1984). These latter estimates are negative (influenced by their prior value of -0.3051, see Table 1) and only adjust towards the long run sample estimate once 40 or more sample observations are incorporated. This result is intuitively reasonable since the counterintuitive negative prior is likely to drive the updated estimates until accumulated sample information takes over. The recursive estimates based on b^*_{02} in (7) are immediately close to the long-run sample estimate and exhibit a remarkable stability relative to the other estimates.

This example illustrates the potential of design reduction. It suggests that dropping the less significant variables in the meta-analysis model can enhance the sampling properties in the cells remaining. This, in turn, will result in accurate prior estimates which in an updating scheme provide stable and efficient recursive parameter estimates. These insightful estimates enable timely model-based decision making. The example also illustrates that the approach does not guarantee improvement, and that the recursive estimates remain susceptible to inherent constraints of Bayesian updating. Nevertheless, the limited empirical results are encouraging and further calibration of the basic approach, perhaps through other updating procedures, is warranted before its full practical value can be established.

4. SUMMARY AND CONCLUSION

Meta-analysis results have been suggested as useful priors in Bayesian estimation in the hope of enhancing parameter efficiency and stability given limited sample sizes. Within a non-probabilistic representation of Bayesian updating, it is shown in this paper that the prior information is confined to the prior estimates which belong to the same cell of the meta-analysis design as the new study for which stable estimates are sought. This raises the problem of empty and scarcely-populated cells inherent in the natural-experiment design underlying the meta-analysis methodology. In deriving prior estimates, analytic discussion and an empirical illustration suggest that reducing the design as to enhance proper sampling in the cells is preferable over regression predictions within a full design.

FOOTNOTES

1 In general, prior estimates could be obtained for all parameters in the model either from separate meta-analyses on each individual parameter (Vanhonacker 1990) or a single meta-analysis on all parameters simultaneously. The latter has not been done in marketing applications but seems preferable because parameter estimates have covariance structures which might be incorporated to enhance parameter efficiency and, hence, sharpen statistical inference.

2 Note that by applying a coding scheme in the design matrix Z such that all entries in z_0 are zeros except the intercept, the biases become less relevant. Specifically, biases in the grand mean (the estimated intercept term in $\hat{c}$) would be the only ones carried over into the prior.

3 Note that the results reported in Table 3 are based on Tellis (1988)'s original coding scheme which is different from the one on which his reported results are based.

APPENDIX 1 : Impact of Design Reduction on Sample Estimates

Consider the column partitioning of full design matrix Z as $Z = (\tilde{Z}_1\ \tilde{Z}_2)$ with the corresponding partitioning of $c' = (c_1'\ c_2')$. It is rather straightforward to show that the GLS estimate $\hat{c}_1$ equals

$$\hat{c}_1 = \left[\tilde{Z}_1' \Delta^{-1} \tilde{Z}_1 - \tilde{Z}_1' \Delta^{-1} \tilde{Z}_2 \left(\tilde{Z}_2' \Delta^{-1} \tilde{Z}_2\right)^{-1} \tilde{Z}_2' \Delta^{-1} \tilde{Z}_1\right]^{-1}$$

$$\left[\tilde{Z}_1 \Delta^{-1} \hat{\beta} - \tilde{Z}_1' \Delta^{-1} \tilde{Z}_2 \left(\tilde{Z}_2' \Delta^{-1} \tilde{Z}_2\right)^{-1} \tilde{Z}_2 \Delta^{-1} \hat{\beta}\right] \qquad (2\text{-}1)$$

with elements as defined in the text.

Under design reduction, the basic model $\hat{\beta} = Zc + \varepsilon$ is estimated as $\hat{\beta} = \tilde{Z}_1 c_1 + v$ (assuming all variables in $\tilde{Z}_2$ are deleted) where $v = \tilde{Z}_2 c_2 + \varepsilon$ with the assumption that $E(v) = 0$.

Accordingly,

$$E(vv') = \Delta + \tilde{Z}_2\, c_2\, c_2'\, \tilde{Z}_2'$$

and the GLS estimate of the reduced design model equals

$$\hat{c}_1^R = \left[\tilde{Z}_1' \left(\Delta + \tilde{Z}_2 c_2 c_2' \tilde{Z}_2'\right)^{-1} \tilde{Z}_1\right]^{-1} \tilde{Z}_1' \left(\Delta + \tilde{Z}_2 c_2 c_2' \tilde{Z}_2'\right)^{-1} \hat{\beta}\ . \qquad (2\text{-}2)$$

Bartlett 's (1951) matrix-inversion identity implies that

$$\left[\Delta + \tilde{Z}_2 c_2 c_2' \tilde{Z}_2'\right]^{-1} = \Delta^{-1} - \frac{1}{\left(1 + c_2' \tilde{Z}_2' \Delta^{-1} \tilde{Z}_2 c_2\right)} \Delta^{-1} \tilde{Z}_2 c_2 c_2' \tilde{Z}_2' \Delta^{-1}.$$

Incorporating this into (2-2), it becomes evident that $\hat{c}_1$ in (2-1) will equal $\hat{c}_1^R$ if and only if the following identity holds

$$\left(\tilde{Z}_2' \Delta^{-1} \tilde{Z}_2\right)^{-1} = \frac{1}{1 + c_2' \tilde{Z}_2' \Delta^{-1} \tilde{Z}_2 c_2} c_2 c_2'\ .$$

This identity holds for either $\left(\tilde{Z}_2' \Delta^{-1} \tilde{Z}_2\right)^{-1} = 0$ or $c_2 = \infty$. As $\left(\tilde{Z}_2' \Delta^{-1} \tilde{Z}_2\right)^{-1}$ is the variance-covariance matrix of the independent c_2 estimates, it appears that as the corresponding t-values (i.e., ratio of parameter estimate over estimated standard error) increase, $\hat{c}_1^R$ will approach $\hat{c}_1$.

Moreover, following the suggested reduction process, as the least significant design variables are dropped first (i.e., the ones with the relatively lowest t-values), the sample estimates of the parameters of the remaining design variables will change but those changes will become smaller and smaller approaching the stopping point.

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Table 1

REPRESENTATION IN TWO META-ANALYSES

	Assmus, Farley and Lehmann (1984)	Tellis (1988)
Sample Size	128	421
Total Number of Cells in Natural Experimental Design	2.6542×10^{6}	1.8208×10^{10}
Number of Cells with at Least One Observation	43	209

Table 2

META-ANALYSIS REGRESSION RESULTS: SHORT-TERM ADVERTISING ELASTICITY

Variable	Original Assmus, Farley, Lehmann (1984) Design		Reduced Design		Design Profile (z_0) New Data Set	
	Parameter Estimates	(Standard Error)	Parameter Estimates	(Standard Error)	(1)[a]	(2)[b]
Intercept	0.3790	(0.1699)	0.3432	(0.0554)	1	1
Dependent	-0.0039	(0.1011)	0.0517	(0.0538)	1	1
Variable	0.1203	(0.1046)	0.0517	(0.0597)	0	0
Advertising	-0.0986	(0.1184)			1	
Variable	-0.0140	(0.0647)			0	
Mature Product	-0.0146	(0.0884)			1	
Carry Over	-0.2471	(0.0747)	-0.1773	(0.0432)	1	1
Other Variables	-0.0338	(0.1027)			-1	
Price	-0.0198	(0.0849)			1	
Exogenous Variables	-0.0791	(0.0542)	-0.0729	(0.0098)	-1	-1
Estimation	-0.1361	(0.0681)	-0.0570	(0.0400)	1	1
Method	0.0622	(0.1044)			0	
	0.0857	(0.1269)			0	
Pooled Data	-0.1386	(0.0785)	-0.0768	(0.0088)	-1	-1
Multiplicative Model	-0.1565	(0.1073)	-0.1242	(0.0390)	1	1
Media Aggregated	-0.1517	(0.1285)	-0.0282	(0.0351)	1	1
TV Advertising	-0.0626	(0.0655)			-1	
Frequently Purchased	0.0230	(0.0619)			1	
Product	-0.0419	(0.0736)			1	
Category	0.0264	(0.0616)			-1	
	-0.0429	(0.1024)			-1	
National	-0.0709	(0.0888)	-0.0741	(0.0512)	1	1
Setting	0.1620	(0.1983)	0.2670	(0.0745)	0	0
Brand Level	0.0305	(0.1078)			1	
Temporal	-0.0226	(0.1215)			0	
Aggregation	-0.1271	(0.0811)			1	

[a] Original (full) design profile.
[b] Reduced design profile.

Table 3

META-ANALYSIS REGRESSION RESULTS : PRICE ELASTICITY

Variable	Original Tellis (1988) Design		Reduced Design		Design Profile (z_0) New Data Set	
	Parameter Estimates	(Standard Error)	Parameter Estimates	(Standard Error)	(1)[a]	(2)[b]
Intercept	-7.3426	(5.2448)	-3.1403	(4.0463)	1	1
Quality	-3.0202	(1.8112)	-2.2420	(1.6740)	0	0
Distribution	-0.5241	(1.5319)			0	
Advertising	2.5397	(1.6258)	2.1453	(1.2220)	1	1
Promotion	2.5069	(2.2994)			0	
Other Variables	-2.2103	(1.3876)	-1.1091	(1.1952)	0	0
Lag Dependent	-0.8036	(1.2693)			1	
Lag Independent	0.6473	(2.1850)			0	
Functional Form	0.1253	(0.4291)			1	
Dependent Variable Definition	0.6413	(0.5170)			0	
Price Variable Definition	0.2862	(0.5046)			0	
Estimation Method	-0.2688	(0.4874)			1	
Durable	3.4431	(1.9884)	3.5850	(1.7761)	0	0
Product Class	-0.7371	(0.3192)	-0.6466	(0.2589)	1	1
Product Life Cycle	-0.1267	(0.9168)			3	
National Setting	-0.3443	(1.1405)			1	
Data Interval	1.0126	(0.5560)	0.5795	(0.3941)	3	3
Data Type	-4.1428	(0.8362)	-3.8226	(0.6797)	2	2
Elasticity	6.9851	(1.8734)	4.8267	(1.4964)	1	1
Data Source	2.6727	(1.2453)	2.3882	(0.9769)	2	2
Data Level	2.0123	(2.6002)			1	

[a] Original (full) design profile.
[b] Reduced design profile.

Table 4

META-ANALYSIS PRIORS

	Short-term Advertising Elasticity (Assmus, Farley, & Lehmann 1984)	Price Elasticity (Tellis 1988)
Regression Predictions for Full Design $(b_{02} = z_0\hat{c})$	-0.3051	1.8835
Corrected Prior $(b^*_{02} = b_{02} - z_0 D\beta_d)$	0.0325	-2.3360

Table 5
ELASTICITY ESTIMATES FOR NEW DATA SET [a]

Number of Sample Observations	Parameter Estimates (Estimated Variance) Based on: Sample Observations[b] — Advertising	Sample Observations[b] — Price	Meta-Analysis Prior b^*_{02} on Both Parameters — Advertising	Meta-Analysis Prior b^*_{02} on Both Parameters — Price
5	0.2310(0.00)	1.7210(7.48)	0.0707(0.01)	-5.3107(0.02)
6	0.1530(0.01)	- 3.4815(12.80)	0.0639(0.00)	-5.3096(0.01)
7	0.1521(0.01)	- 3.4374(5.39)	0.0666(0.00)	-5.3081(0.01)
8	0.1544(0.00)	- 3.4914(4.67)	0.0672(0.00)	-5.3032(0.02)
9	0.1549(0.00)	- 4.0184(4.70)	0.0662(0.00)	-5.3030(0.02)
10	0.1520(0.00)	- 4.1684(4.05)	0.0650(0.00)	-5.3067(0.02)
11	0.0950(0.01)	- 3.5544(6.43)	0.0511(0.00)	-5.3080(0.02)
12	0.0462(0.01)	- 1.6317(10.70)	0.0383(0.00)	-5.3088(0.02)
13	0.0457(0.01)	- 3.5097(11.00)	0.0351(0.00)	-5.3059(0.02)
14	0.0554(0.01)	- 3.5970(9.87)	0.0390(0.00)	-5.3055(0.02)
15	0.0635(0.01)	- 4.2443(7.81)	0.0422(0.00)	-5.3032(0.02)
16	0.0582(0.01)	- 1.3461(5.05)	0.0431(0.00)	-5.3026(0.02)
17	0.0645(0.01)	- 1.0044(4.21)	0.0475(0.00	-5.3002(0.02)
18	0.0548(0.00)	- 1.0553(3.78)	0.0558(0.00)	-5.3000(0.02)
19	0.0566(0.00)	- 1.0573(3.53)	0.0592(0.00)	-5.3006(0.02)
20	0.0610(0.00)	- 1.2296(3.26)	0.0605(0.00)	-5.3041(0.00)

[a] Multiplicative sales response model with predictors advertising share, relative price, and lagged market share.

[b] Ordinary Least Squares (OLS) estimates.

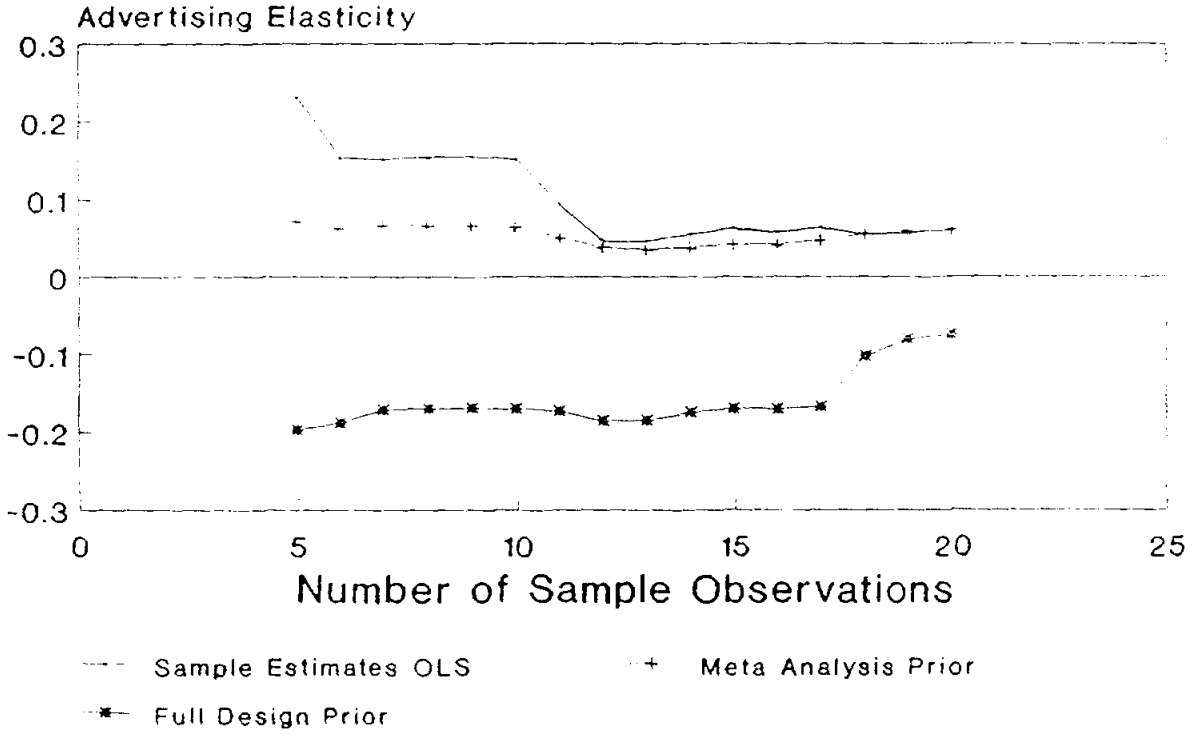

Figure 1

ADVERTISING ELASTICITY ESTIMATES

FOR NEW DATA SET

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