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BEFORE THE
POSTAL RATE COMMISSION
WASHINGTON, D. C. 20268-0001

POSTAL RATE AND FEE CHANGES, 2000	Docket No. R2000–1

RESPONSE
OF
A. THOMAS BOZZO
TO NOTICE OF INQUIRY NO. 4
ON BEHALF OF THE
UNITED STATES POSTAL SERVICE

August 21, 2000

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Purpose and Scope

The purpose of this document is to provide econometric estimates responsive to item (a) in the Commission's Notice of Inquiry ("NOI") No. 4 Concerning Mail Processing Variability Models. The text of the NOI is attached. Dr. Greene's response addresses the theoretical issues raised in items (b)-(f) in the NOI. I also provide econometric variability estimates based on the specification that Dr. Greene calls "Model C", which is the general model in which the NOI's "Model A" and "Model B" are nested. Supporting materials are provided in USPS–LR–I–461.

I. Introduction

The NOI asks parties to "test the compatibility of witness Bozzo's data with the family of models that lack facility-indexed coefficients." First, I note that I performed exactly such a test to validate my choice of the panel data fixed effects estimator for the results I recommend in my direct testimony. USPS–T–15 at 122-124. The results of the standard specification tests I performed unambiguously reject the "pooled" model, which lacks facility specific coefficients, as well as the model with random facility effects. The implication is that the estimates from the pooled, between, and random effects models, all of which incorporate statistical restrictions that are rejected per the specification tests, are biased and inconsistent. Accordingly, the evidence on the record of this proceeding shows the pooled model (hereafter "Model 0"), without facility-indexed coefficients, to be rejected in favor of the fixed-effects estimator of what the NOI terms "Model A," with facility-indexed intercepts.

The NOI also defines "Model B," which is a panel data model with time-indexed but not facility-indexed intercepts. Item (a) in the NOI requests that parties test whether Model 0 (the pooled model) can be rejected in favor of Model B, and whether a fixed- or random-effects formulation is appropriate to estimate Model B, for the five largest MODS operation groups covered by my analysis.[1] In Section II, I present the results of specification tests responsive to item (a) of the NOI. In his response to the NOI, Dr. Greene notes that specification tests comparing Model B with Model 0 have no bearing on the fundamental issue of whether there are significant facility-specific effects.

Dr. Greene indicates that Model A and Model B are not nested, but that both Model A and Model B are nested in what he terms "Model C." Dr. Greene's Model C incorporates both time-specific and facility-specific effects. The questions implied by the NOI are whether adding time effects to Model A, or *facility effects to Model B, materially changes the results. In Section III, I present* results that address these questions. The results show that, taking Model B as the starting point, it is possible to decisively reject Model B in favor of Model C with both facility- and time-indexed intercepts. The result is analogous to the rejection of Model 0 in favor of Model A. Also In Section III, I compare the results

[1] The five largest MODS operation groups, by cost pool dollars, are (in alphabetical order), BCS, FSM, Manual Flats, Manual Letters, and SPBS. Below, I also present results for the OCR operation group, since my TSP programs produce results for the mechanized and automated letter and flat operations as a group.

I obtained in my direct testimony (Model A) with those of Model C. I show that the variabilities resulting from Model C are essentially the same as those that I present in my direct testimony.

II. Econometric specification of "Model B", and results of the specification tests described in part (a) of the NOI

In this section, I present the econometric specification of the pooled Model 0 and of Model B. I also present the results of specification tests of model B against Model 0. The specification of the pooled Model 0 that served as the basis for the tests of Model A in my testimony is:

$$
\begin{aligned}
\ln HRS_{it} = {} & \beta_0 + (\alpha_1 + \gamma_1 L + \gamma_2 L^2 + \gamma_3 L^3 + \gamma_4 L^4)\ln TPH_{it} \\
& +(\alpha_{11} + \gamma_{11} L + \gamma_{22} L^2 + \gamma_{33} L^3 + \gamma_{44} L^4)(\ln TPH_{it})^2 \\
& +\alpha_2 \ln CAP_{it} + \alpha_{22}(\ln CAP_{it})^2 + \alpha_3 \ln DEL_{it} + \alpha_{33}(\ln DEL_{it})^2 \\
& +\alpha_4 \ln WAGE_{it} + \alpha_{44}(\ln WAGE_{it})^2 + \alpha_5 TREND_{it} + \alpha_{55} TREND_{it}^{\,2} \\
& +\alpha_6 \ln MANR_{it} + \alpha_{66}(\ln MANR_{it})^2 \\
& +\alpha_{12} \ln TPH_{it} \ln CAP_{it} + \alpha_{13} \ln TPH_{it} \ln DEL_{it} + \alpha_{14} \ln TPH_{it} \ln WAGE_{it} \\
& +\alpha_{15} \ln TPH_{it} \cdot TREND_{it} + \alpha_{16} \ln TPH_{it} \ln MANR_{it} \\
& +\alpha_{23} \ln CAP_{it} \ln DEL_{it} + \alpha_{24} \ln CAP_{it} \ln WAGE_{it} + \alpha_{25} \ln CAP_{it} \cdot TREND_{it} \\
& +\alpha_{26} \ln CAP_{it} \ln MANR_{it} \\
& +\alpha_{34} \ln DEL_{it} \ln WAGE_{it} + \alpha_{35} \ln DEL_{it} \cdot TREND_{it} \\
& +\alpha_{36} \ln DEL_{it} \ln MANR_{it} \\
& +\alpha_{45} \ln WAGE_{it} \cdot TREND_{it} + \alpha_{46} \ln WAGE_{it} \ln MANR_{it} \\
& +\alpha_{56} TREND_{it} \ln MANR_{it} \\
& +\beta_2 QTR2_{it} + \beta_3 QTR3_{it} + \beta_4 QTR4_{it} \\
& +\varepsilon_{it}.
\end{aligned}
$$

Model 0 differs from the specification of Model A (see USPS–T–15 at 117) in that the intercept term is assumed not to vary with the facility, indexed by i.[2] Note that it contains trend terms and seasonal (quarterly) dummy variables, which would be expected to capture some (if not most) of the time-specific effects specified in Model B. The Model 0 equation given above is applicable to the letter and flat shape operations. The corresponding SPBS equation omits terms involving the manual ratio variable. See USPS–T–15 at page 118. The corresponding estimating equation for Model B is:

$$
\begin{aligned}
\ln HRS_{it} = {} & \beta_0 + \lambda_t + (\alpha_1 + \gamma_1 L + \gamma_2 L^2 + \gamma_3 L^3 + \gamma_4 L^4)\ln TPH_{it} \\
& + (\alpha_{11} + \gamma_{11} L + \gamma_{22} L^2 + \gamma_{33} L^3 + \gamma_{44} L^4)(\ln TPH_{it})^2 \\
& + \alpha_2 \ln CAP_{it} + \alpha_{22}(\ln CAP_{it})^2 + \alpha_3 \ln DEL_{it} + \alpha_{33}(\ln DEL_{it})^2 \\
& + \alpha_4 \ln WAGE_{it} + \alpha_{44}(\ln WAGE_{it})^2 + \alpha_5 TREND_{it} + \alpha_{55} TREND_{it}^{\,2} \\
& + \alpha_6 \ln MANR_{it} + \alpha_{66}(\ln MANR_{it})^2 \\
& + \alpha_{12} \ln TPH_{it} \ln CAP_{it} + \alpha_{13} \ln TPH_{it} \ln DEL_{it} + \alpha_{14} \ln TPH_{it} \ln WAGE_{it} \\
& + \alpha_{15} \ln TPH_{it} \cdot TREND_{it} + \alpha_{16} \ln TPH_{it} \ln MANR_{it} \\
& + \alpha_{23} \ln CAP_{it} \ln DEL_{it} + \alpha_{24} \ln CAP_{it} \ln WAGE_{it} + \alpha_{25} \ln CAP_{it} \cdot TREND_{it} \\
& + \alpha_{26} \ln CAP_{it} \ln MANR_{it} \\
& + \alpha_{34} \ln DEL_{it} \ln WAGE_{it} + \alpha_{35} \ln DEL_{it} \cdot TREND_{it} \\
& + \alpha_{36} \ln DEL_{it} \ln MANR_{it} \\
& + \alpha_{45} \ln WAGE_{it} \cdot TREND_{it} + \alpha_{46} \ln WAGE_{it} \ln MANR_{it} \\
& + \alpha_{56} TREND_{it} \ln MANR_{it} \\
& + \beta_2 QTR2_{it} + \beta_3 QTR3_{it} + \beta_4 QTR4_{it} \\
& + \varepsilon_{it}.
\end{aligned}
$$

The terms $\beta_0 + \lambda_t$ in the equation above correspond to the term α_t in the NOI. The NOI notes that regressors made redundant by the inclusion of time effects in Model B may be omitted. None of the other regressors should be

[2] I also relabled the intercept term to be consistent with the notation in Dr. Greene's response.

excluded from the pooled model estimated for the specification tests. Since the relevant issue is whether the trend and seasonal variables in Model 0 adequately control for time-specific effects, it is not appropriate to exclude any variables from Model B *a priori*. In order to avoid the significant programming complications that would be required to apply an appropriate autocorrelation adjustment to the random effects estimator for Model B, I estimated Model 0 and Model B without the autocorrelation adjustment.[3] These results are given in Table 1.

The NOI requests in part (a) that respondents test (1) the null hypothesis of a common intercept for all time periods (" $\alpha_t = \alpha$ for all t") against the alternative that the intercepts vary over time—i.e., Model 0 versus Model B— and (2) random effects versus fixed effects applied to Model B. These test are appropriately conducted by using an F statistic and a Hausman test, respectively. Both test statistics are computed by the programs named var(ltr,nl)-(tpf,tph)-by98-noi4b.tsp, in LR–I–461. The test results are presented in Table 1.

The P-values of the specification test statistics are reported in Table 1. The test statistic values and degrees of freedom are presented in the regression output in LR–I–461. The F test for common intercepts over time indicates that Model 0 cannot be rejected in favor of Model B for four of the six operation groups I examined: OCR, SPBS, FSM, and Manual Flats. However, these results do not weigh in favor of Model 0 since, for those cost pools, Model 0 has already been rejected in favor of Model A using the specification tests reported in

[3] Failing to adjust for autocorrelated disturbances impacts the efficiency, but not the unbiasedness and consistency, of the estimates (see, e.g., William H. Greene, *Econometric Analysis*, Second Edition, at 418-419).

USPS–T–15 (see USPS-T-15 at 122-124). One other case, BCS, is "borderline"—Model 0 is rejected at the 5 percent significance level but not at the one percent significance level. Only Manual Letters shows strong evidence in favor of Model B over Model 0.

The Hausman test indicates that the random effects model cannot be rejected in favor of the fixed effects model in any of the six cost pools. The results should not be surprising. The pooled model already contains a quadratic trend term and seasonal dummy variables. To the extent that those variables are successful at capturing the period-specific effects, the time-indexed intercepts in Model B should not add much explanatory power to the model. Nor does the presence or absence of period-specific effects say anything about whether facility-indexed intercept components also belong in the model, as explained by Dr. Greene in his response to this NOI. To address that issue, it is necessary to estimate Dr. Greene's Model C, which I do in the next section.

Table 1.
Specification test results for "Model B"

Cost Pool	BCS	OCR	FSM	SPBS	Manual Flats	Manual Letters
P-value, F test: H_0: Model 0 (with no time effects) vs. H_1: Model B (with time effects)	0.0150	0.5458	0.1550	0.5202	0.9739	0.0006
P-value, Hausman test: H_0: Model B (fixed effects) vs. H_1: Model B (random effects)	0.4179	>0.9995*	0.6132	>0.9995*	>0.9995*	0.3393
Reject H_0: Model 0 (with no time effects) vs. H_1: Model B (with time effects)?	Borderline**	No	No	No	No	Yes
Reject H_0: Model B (fixed effects) vs. H_1: Model B (random effects)?	No	No	No	No	No	No

*P-value is 1 to all reported digits.
**H_0 is rejected at the 5 percent significance level, but not the 1 percent significance level.

III. Econometric specification and estimates of "Model C"

Both Model A and Model B can be represented as special cases of Model C, which includes both facility-indexed and time-indexed components in the regression intercept. The estimating equation for Model C is:

$$
\begin{aligned}
\ln HRS_{it} = {} & \beta_0 + \delta_i + \lambda_t + (\alpha_1 + \gamma_1 L + \gamma_2 L^2 + \gamma_3 L^3 + \gamma_4 L^4)\ln TPH_{it} \\
& + (\alpha_{11} + \gamma_{11} L + \gamma_{22} L^2 + \gamma_{33} L^3 + \gamma_{44} L^4)(\ln TPH_{it})^2 \\
& + \alpha_2 \ln CAP_{it} + \alpha_{22}(\ln CAP_{it})^2 + \alpha_3 \ln DEL_{it} + \alpha_{33}(\ln DEL_{it})^2 \\
& + \alpha_4 \ln WAGE_{it} + \alpha_{44}(\ln WAGE_{it})^2 + \alpha_5 TREND_{it} + \alpha_{55} TREND_{it}^2 \\
& + \alpha_6 \ln MANR_{it} + \alpha_{66}(\ln MANR_{it})^2 \\
& + \alpha_{12} \ln TPH_{it} \ln CAP_{it} + \alpha_{13} \ln TPH_{it} \ln DEL_{it} + \alpha_{14} \ln TPH_{it} \ln WAGE_{it} \\
& + \alpha_{15} \ln TPH_{it} \cdot TREND_{it} + \alpha_{16} \ln TPH_{it} \ln MANR_{it} \\
& + \alpha_{23} \ln CAP_{it} \ln DEL_{it} + \alpha_{24} \ln CAP_{it} \ln WAGE_{it} + \alpha_{25} \ln CAP_{it} \cdot TREND_{it} \\
& + \alpha_{26} \ln CAP_{it} \ln MANR_{it} \\
& + \alpha_{34} \ln DEL_{it} \ln WAGE_{it} + \alpha_{35} \ln DEL_{it} \cdot TREND_{it} \\
& + \alpha_{36} \ln DEL_{it} \ln MANR_{it} \\
& + \alpha_{45} \ln WAGE_{it} \cdot TREND_{it} + \alpha_{46} \ln WAGE_{it} \ln MANR_{it} \\
& + \alpha_{56} TREND_{it} \ln MANR_{it} \\
& + \beta_2 QTR2_{it} + \beta_3 QTR3_{it} + \beta_4 QTR4_{it} \\
& + \varepsilon_{it}.
\end{aligned}
$$

I present econometric variability estimates for Model C, as well as the results of specification tests of Model C against Model B, in Table 2. The results closely mirror those from Model A that I present in my direct testimony (USPS–T–15 at 119-120. The variabilities presented in Table 2 are based on the fixed effects estimates of Model C, and adjust for autocorrelation of the disturbances. The F-test, which here tests Model B (without facility-indexed intercepts) against Model C (with facility-indexed intercepts), strongly rejects Model B in favor of Model C for all six cost pools. Furthermore, the Hausman test of random effects versus fixed effects for Model C supports the fixed-effects model over the random effects model.

Table 2.
Principal results and specification test statistics for "Model C"

Cost Pool:	BCS	OCR	FSM	SPBS	Manual Flats	Manual Letters
Output Elasticity (Volume-variability factor)*	0.877 (0.030)	0.742 (0.039)	0.840 (0.026)	0.664 (0.045)	0.764 (0.028)	0.732 (0.025)
Auto-correlation *coefficient*	0.643	0.701	0.627	0.594	0.673	0.699
Adjusted R-squared	0.986	0.972	0.994	0.987	0.988	0.991
Number of observations	5,406	5,097	4,373	1,584	4,891	5,512
Number of sites	298	289	236	95	278	300
P-value, F test: H_0: Model B (with no facility effects) vs. H_1: Model C (with facility effects)	**	**	**	**	**	**
P-value, Hausman test: H_0: Model C (fixed effects) vs. H_1: Model C (random effects)	0.0022	**	0.0007	0.0011	**	**
Reject H_0: Model B (with no facility effects) vs. H_1: Model C (with facility effects)?	Yes	Yes	Yes	Yes	Yes	Yes
Reject H_0: *Model* C (with fixed effects) vs. H_1: *Model C (with* random effects)	Yes	Yes	Yes	Yes	Yes	Yes

*Elasticities evaluated using full data set and arithmetic mean method; standard errors in parentheses.
**<0.00005 (P-value is 0 to all reported digits).

I compare the estimated variabilities from Model A and Model C in Table 3. The results from Model A and Model C are very similar, and the results for individual cost pools differ by less than the estimated standard errors of the variability estimates. There is no indication of systematic bias, and the composite variability for the six cost pools examined here differs by only 0.1 percent between Model A and Model C. The result is consistent with the finding above that the time-specific intercepts contribute little additional information over the trend and quarterly variables for most cost pools. Accordingly, the results in USPS–T–15 are not "fragile" when compared to those of the more general Model C.

Table 3.
Comparison of "Model A" and "Model C" variabilities

Cost Pool	"Model A" Variability (USPS–T–15 at 119-120)	"Model C" Variability (LR–I–461)	Percentage difference: "Model A" vs. "Model C"
BCS	0.895	0.877	-2.0%
OCR	0.751	0.742	-1.2%
Manual Flats	0.772	0.764	-1.0%
Manual Letters	0.735	0.732	-0.4%
FSM	0.817	0.840	2.8%
SPBS	0.641	0.664	3.6%
Composite	0.786	0.787	0.1%

IV. Summary

In this analysis, I demonstrate that the specification defined in the NOI as Model B (a panel data model with time-indexed by not facility-indexed intercepts) generally adds little explanatory ability compared to the pooled model already presented and rejected in USPS-T-15. Furthermore, Model B can be decisively rejected in favor of a more general specification, Model C. Model C, as discussed by Dr. Greene in his response to this NOI, incorporates both time-specific and facility-specific effects. The available evidence—both statistical, as presented here and in USPS-T-15, and operational, as described by witness Degen at pages 18-23 of USPS-T-16—overwhelmingly supports the existence of facility-specific, non-volume factors that affect costs. The rejected models, the pooled and "between" models as well as Model B, inappropriately ignore the facility-specific effects and are seriously biased. Since the biases of the rejected models have no relevant economic interpretation, but simply reflect a confounding of volume and non-volume factors, they do not provide reliable estimates of volume-variability factors for mail processing and should not be adopted. My analysis also shows that the results I present in my direct testimony (for Model A) are essentially the same as those obtained from Model C. The results I present in USPS-T-15 are robust to the inclusion of the period-specific effects that yield the more general Model C, which provides further evidence that the USPS-T-15 results are reliable and should be adopted.

REPRODUCTION OF THE TEXT OF:

NOTICE OF INQUIRY NO. 4

CONCERNING MAIL PROCESSING VARIABILITY MODELS

(Issued August 2, 2000)

In Docket No. R97-1, witness Bradley conducted a specification search for a model of mail processing variability. He tested a family of models that lack time-indexed coefficients, and rejected the more restrictive models in favor of the facility-specific fixed-effects model. In response to Notice of Inquiry No. 4 in R97-1, the facility-specific fixed-effect model was tested and rejected against the general model, which had both time-indexed and facility-indexed coefficients. In Docket No. R97-1, witness Neels commented that this specification search had produced "too fragile and incomplete a set of results." One respect in which Mr. Neels regarded Dr. Bradley's specification search as incomplete was its failure to evaluate a parallel family of models that lacks facility-specific coefficients. See Docket No. R97-1 at Tr. 28/15775-84, 15805. This family of models was described in Docket No. R97-1 at Tr. 15776.

The record in this docket appears to be incomplete in the same respect as the record in Docket No. R97-1. To help provide a more complete record in this docket, interested parties are invited to test the compatibility of witness Bozzo's data with the family of models that lack facility-indexed coefficients. They are also invited to discuss, in testimony or comments, whether these specification test results, or those already performed by witness Bozzo, establish the validity of any particular model or family of models. Responses are due within 14 days of the date of this Notice.

Specifically, interested parties are invited to consider the model tested by witness Bozzo that lacks time-indexed coefficients. It will be labeled Model A and it takes the general form

$$y_{it} = \alpha_i + x_{it}\,\beta + \varepsilon_{it}$$

Here α_i denotes a facility-specific fixed-effect, y_{it} is the logarithm of hours in that operation, and x_{it} is the vector of variables including the logarithm of total piece-handling. Interested parties are also invited to consider an alternative model labeled, Model B,which lacks facility-indexed coefficients. It takes the form

$$y_{it} = \alpha_t + x_{it}\,\beta + \varepsilon_{it}$$

where α_t denotes a quarter-specific fixed effect, and all other variables are as defined above. In both of these models, the subscript i denotes facilities, and the subscript t denotes quarters.

a) Witness Bozzo performs a statistical test of the null hypothesis that $\alpha_i = \alpha$ for all i and rejects this null hypothesis. In addition, he tests and rejects the null hypothesis that the α_i are independently, identically distributed random variables with mean zero and variance. He uses both of these hypothesis tests to demonstrate that the facility-specific fixed effect model is statistically superior to the models nested within it, such as the "pooled" and "random effects" models. For the five largest MODS pools modeled by witness Bozzo (in terms of accrued costs), parties are asked to use his data to perform the following two hypothesis tests with respect to Model B: 1) the null hypothesis that $\alpha_t = \alpha$ for all t , and the null hypothesis that the α_t are independently, identically distributed random variables with mean zero and variance σ^2. Any terms used by witness Bozzo that are not needed because of the presence of α_t , such as lagged dependent variables and regressors may be omitted.

b) Parties are asked to indicate whether rejection of the hypotheses described in a) establish that Model A is statistically superior to the models nested within it, such as the "pooled" and the "random effects" models. Similarly, parties are asked to indicate whether rejection of the hypotheses described in a) establish that Model B is statistically superior to the models nested within it, such as the "pooled" and the "random effects" models.

c) Parties asked to discuss whether Models A and B are nested within one another, and whether rejection of the hypotheses described in a) provide statistical grounds for preferring either of these models over the other.

d) Parties are asked to discuss whether witness Bozzo's rejection of the hypotheses applicable to Model A is sufficient to establish that Model (A) yields a valid estimate of β, which determines the magnitude of volume variability.

e) Parties are asked to discuss whether rejection of the hypotheses applicable to Model (B) is sufficient to establish that Model B yields a valid estimate of β, which determines the magnitude of volume variability.

f) Parties are asked to discuss whether, even with the rejection of the hypotheses described in a), there may be theoretical grounds for concluding that a rejected model could provide a better estimate of variability than either model A or B.

DECLARATION

I, A. Thomas Bozzo, declare under penalty of perjury that the foregoing answers are true and correct, to the best of my knowledge, information, and belief.

A. Thomas Bozzo

Dated: 8/21/00