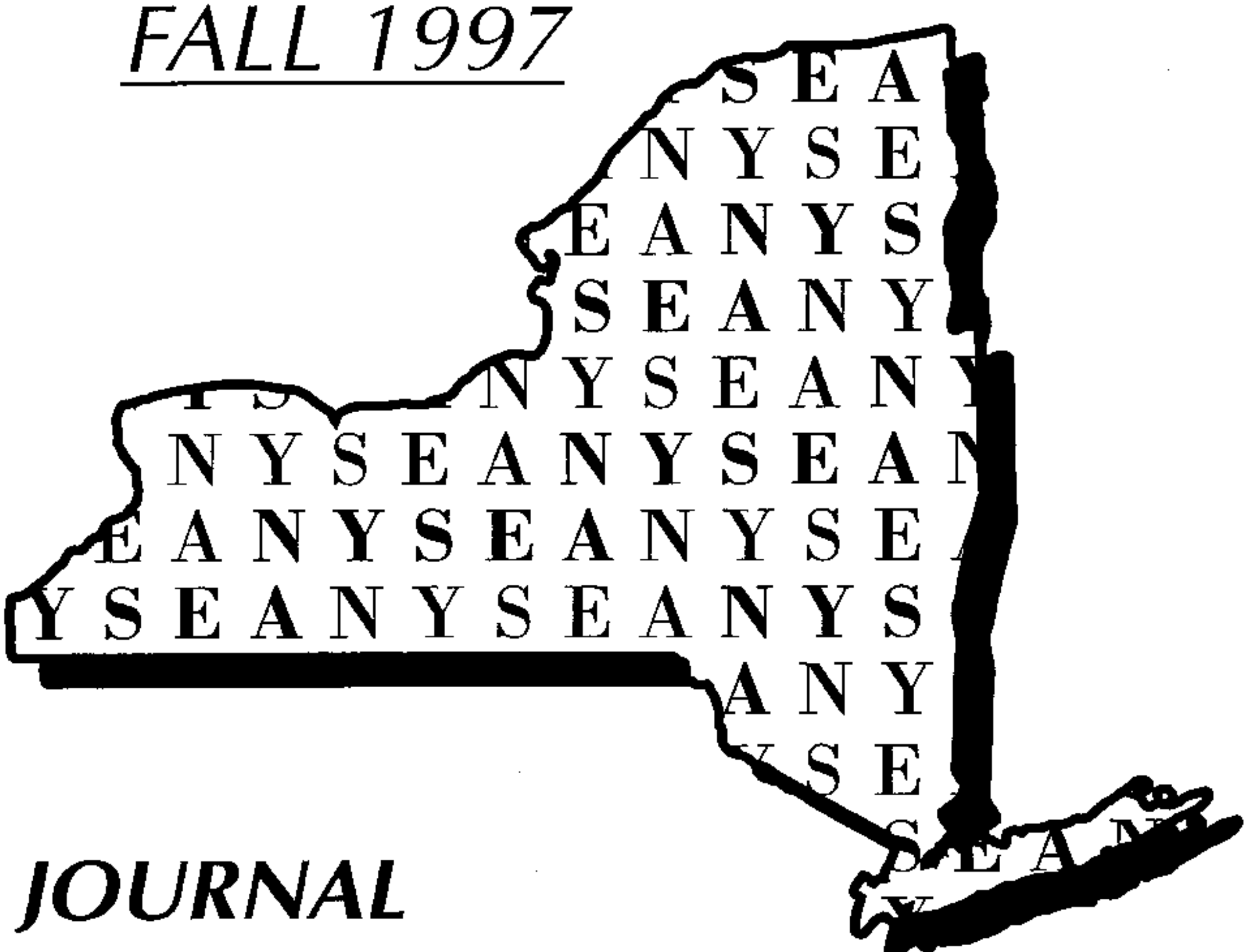
NEW YORK ECONOMIC REVIEW

FALL 1997



JOURNAL of the NEW YORK STATE ECONOMICS ASSOCIATION

VOLUME XXVIII

New York Economic Review

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ARTICLES

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THE EFFECT OF BRANCH VERSUS UNIT BANKING ON SMALL RURAL BUSINESSES IN IDAHO AND MONTANA

Robert J. Tokle* and Mark K. McBeth**

INTRODUCTION

Over the past few years states have established rural development councils and various rural development policies in the hope of improving rural economies. States and their communities, however, are greatly affected by laws passed by the Federal government, such as the Riegle-Neal Interstate Banking and Branching Efficiency Act of 1994. State policy makers require some understanding of the potential impact of the act upon banking and rural economic activity in their states.

Banking organizations can expand geographically by two means: by branching or by a bank holding company buying controlling interest in other banks. Branch banking is where a head office opens up a branch office at a different location. Prior to 1994, under the McFadden Act of 1927, branching across state lines was prohibited. The Riegle-Neal Interstate Banking and Branching Efficiency Act of 1994, however, will allow for nationwide branch banking by June 1, 1997.[1] At the state level, until the Riegle-Neal Act takes effect, it has been up to individual states to decide whether to allow branching.

A bank holding company (BHC) owns controlling interest in one or more banks and often owns other financially related firms. In most states, a BHC can expand geographically by acquiring a controlling interest in other banks. However, the Bank Holding Act of 1956 gave states the right to restrict out-of-state BHCs from acquiring banks in their jurisdictions. As a result, there were very few interstate acquisitions of banks by BHCs until 1975, when Maine allowed BHCs from the New England area to acquire its banks. Maine changed its law because it hoped to attract more bank capital. Since 1975, most states have removed many of the restrictions on acquisitions of banks by out-of-state BHCs, resulting in a large increase in interstate acquisitions during the 1980s and 1990s.

Rural communities and rural small businesses are concerned that bank expansion, either via branching or BHC acquisitions,[2] provides these banks with more monopoly power, thereby, allowing them to make fewer loans locally, charge higher interest rates for loans, and pay lower interest rates on deposits. After a brief review of the literature on the effects of both branch banking and BHC activity on rural communities,[3] we analyze 1993 survey data from small businesses asked about their banking markets in rural Idaho (a branch banking state) and rural Montana (then primarily a unit banking state). This literature and data should present state officials with some understanding of the potential impact of the new law upon rural communities in their states.

LITERATURE SUPPORTING BRANCH BANKING IN RURAL COMMUNITIES

With respect to a number of issues, the literature suggests that branch banking may help or at least not harm rural areas. One issue is competition. McCall (1980), in a highly regarded survey on bank structure and local services, suggests that branch banking increases the number of banks and bank offices in rural areas (p. 103). This implies lower bank concentration and more bank competition in

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branch-banking rural communities. He wrote (p. 108) that both rural and urban banks in branch banking states paid higher interest rates on deposits and had lower profits than their unit banking state counterparts.

Another issue is whether branch banks might take in local deposits and use them to make loans and other investments outside the community. This "siphoning-off" of local deposits is a concern in rural areas. However, the evidence from a number of studies suggests that this is not a problem. In his survey article, McCall (p.104), found that banks and multi-BHCs consistently had a larger percentage of their assets in loans than unit banks.[4] Since it is usually difficult, however, to gather data on what percentage of the loans made by branch banks or multi-BHCs are made to local borrowers, he wrote that the literature "does not substantiate (or refute)" this concern (p. 104).

Barkley, et al. (1984), however, found unique data to study this issue. The data for 1977 to 1980 came from banks in Arizona (a branch banking state) and Colorado (then, a unit banking state). They had access to data on deposits and loans from individual branches where each branch "reported loans and deposits for the specific community involved and cross-office loan participation was represented" (p. 285). The data demonstrated that while branch banks in Arizona might transfer loanable funds from rural to urban areas, the actual local loan-to-deposit ratios were very similar (Colorado .693 and Arizona .685) and not significantly different between the rural banks in each state. Barkley, et al. concluded that as branch banking increases, rural communities, as a whole, will not suffer from less credit.

In addition, Green (1986), Markley (1987) and Lawrence and Klugman (1991) all found that in rural areas, BHC-affiliated banks had higher loan-to-deposit ratios than did independent banks. Also, Barrett and Unger (1991) found that for Montana, BHC affiliates had higher loan-to-asset ratios (a measure similar to loan-to-deposits) than did unit banks. Furthermore, McCall (1980) noted that since unit banks operate in a single geographical area, they reduce their portfolio risk by making fewer loans.

With respect to making agricultural and small business loans, McCall (1980, p. 106) found no discrimination by branch banks or multi-BHCs. Also, studies have found that branch banks are able to offer a greater range of services in rural areas than the smaller unit banks can (McCall 1980, p. 104 and Milklove, 1985). And, since branch banks are more diversified, they are more stable (Smith, 1987). Finally, there probably exists an economies of scale advantage in branch banking (Milkove, 1985).

LITERATURE OPPOSED TO BRANCH BANKING IN RURAL COMMUNITIES

However, there is theory and evidence suggesting that branch banking is harmful to rural communities. Struck and Mandell (1983, p. 1030) conclude "that the credit needs of small businesses are less likely to be met in states that permit wider geographic activities, i.e., statewide and limited branching, and unit banking with multi-bank holding company provisions. This suggests that the level of local decision making as influenced by the banking structure may be of paramount importance." Struck and Mandell and Milkove (1985), argue that in unit banking states, decisions are made by local loan officers and therefore may take into account the character of the borrower in addition to financial information. Conversely, in branch banking states, loan decisions are generally made by the central office. The results of Struck and Mandell can not conclusively be applied to rural communities, however, since their results apply to a survey of both urban and rural small businesses.

Some newspaper accounts also favor unit banking over branch banking for small rural businesses. Motsch (1988, p. 6) in the *American Banker,* quotes sociology professor Robert Ratcliffe:

"When a bank has been taken over and all of the loans are transferred to big-city headquarters, the chances of getting an unusual loan may be zero." Apcar (1988) in the *Wall Street Journal* wrote about deposits being siphoned out of small towns in Texas. Deposits generated in small towns were used to refinance real-estate and energy loans in Dallas and Houston while small town businesses suddenly had their lines of credit cut or notes recalled for no apparent reason.

The majority of the literature suggests that branch banking and BHC organizations will not harm and may even be helpful to rural communities. Struck and Mandell (1983) and some newspaper accounts, however, question this conclusion. To test the question of whether branch banking is beneficial for rural communities, we analyze survey data from small businesses in rural Idaho and Montana. Unlike the nationwide (rural and urban) survey used by Struck and Mandell (1983), our survey covered only small rural businesses.

DATA

Surveys were mailed to 100 businesses in Idaho and 100 businesses in Montana.[5] The businesses were selected by first identifying communities in each state with a population under 10,000. Five communities from each state were then randomly selected from these identified small communities. Businesses were identified from each of these ten communities through community chamber of commerce lists. Surveys were mailed to twenty businesses randomly selected from each community's chamber list. Budgetary constraints prevented a follow up mailing. Forty-five surveys were returned from Montana and 38 from Idaho for an overall response rate of 42 percent. A profile of respondents shows that 85 percent of responding businesses in Montana and 79 percent in Idaho were businesses with less than 20 employees. Table 1 presents a profile of respondents.

TABLE 1. PROFILE OF RESPONDENTS

Montana	Idaho
Number of Employees:	**Number of Employees:**
Full Time: Mean = 8.2	Full Time: Mean = 13.9
Part Time: Mean = 4.0	Part Time: Mean = 4.40
Total Employees: Mean = 12.2	Total Employees: Mean = 18.30
n = 45	n = 38

HYPOTHESES

Hypothesis #1: "We would expect that branch banks charge lower rates of interest on loans than unit banks." The hypothesis is based on the following literature. McCall (1980) found that branch banking increases the number of banks and bank offices in rural areas, leading to increased competition and lower interest rates on loans. Barrett and Unger (1991) found that as the size of the bank increases the interest rate charged on loans decreases because of economies of scale. Since unit banks are smaller, we would expect them to charge higher rates of interest on loans.

Hypothesis #2: "We would expect that branch banks pay higher interest rates on checking accounts, savings accounts, certificates of deposits, and money market accounts than unit banks." McCall's argument (1980) that branch banking leads to increased competition leads to this second hypothesis.

Hypothesis #3: "We would expect no association between customer satisfaction with their bank's willingness to make loans and the type of banking in the state." McCall (1980) found there is no evidence that branch banking has transferred funds from rural to metropolitan areas and there is no evidence of discrimination against agriculture or small businesses. Barkley, et al. (1984) found no differences in loan-to-deposit ratios between branch and unit banking. In addition, Green (1986), Markley (1987) and Lawrence and Klugman (1991) found that affiliate banks of BHC's had higher loan-to-deposit ratios.

Hypothesis #4: "We would expect that respondents in the unit banking state would rate their banks' understanding of the local economy more highly than respondents in the branch banking state." Struck and Mandell (1983) and Milkove (1985) argue that in unit banking states, decisions are made locally.

RESULTS

Hypothesis #1: Interest Rates on Loans

Respondents reported that banks in Idaho charged an average interest rate of 8.28 percent compared to 9.41 percent in Montana. This was a statistically significant difference (t = 2.017, df =37, one tailed test) (see Table 2). We therefore reject the null hypothesis that there is no difference in interest rates on loans between banks in Idaho and Montana.

TABLE 2. INTEREST RATES CHARGED ON LOANS[a]

	Montana	Idaho	t test	df
Respondents	45	38		
Interest Rate on Loans	9.41% (N=26) S.D. = 1.58	8.28% (N=13) S.D. = 1.48	2.017*	37

[a] The loans reported were short-term (one year or less).
* Significant at .05 level.

Hypothesis #2: Interest Paid

There were no statistically significant differences between the interest rates paid on checking, savings, and cd's. There was, however, a significant difference between interest rates paid on money markets accounts (t = 2.33, df =17, one-tailed) with Idaho banks paying the higher interest rate (see Table 3).

TABLE 3. INTEREST RATES PAID ON CHECKING, SAVING, CD'S, AND MONEY MARKET ACCOUNTS

Checking					
	N	**Mean**	**S.D.**	**Z Test**	
Idaho	38	1.43%	1.29	0.16	
Montana	45	1.37%	1.43		
Savings					
	N	**Mean**	**S.D.**	**t test**	**df**
Idaho	23	3.32%	0.72	0.25	44
Montana	23	3.29%	0.73		
C.D's					
	N	**Mean**	**S.D.**	**t test**	**df**
Idaho	9	3.70%	0.72	-1.08	25
Montana	18	3.73%	0.74		
Money Market					
	N	**Mean**	**S.D.**	**t test**	**df**
Idaho	7	4.61%	1.63	2.33*	17
Montana	12	3.98%	1.83		

* Statistically significant at .05.

Hypothesis #3: Satisfaction with Bank's Loan Policies

Seventy-eight percent of Montana respondents were satisfied with their local bank's willingness to make loans compared to 70 percent in Idaho. Despite this difference, the chi-square of 0.70 revealed an insignificant association (see Table 4).

TABLE 4. SATISFACTION WITH BANK'S WILLINGNESS TO MAKE LOANS

"How satisfied are you with your bank's willingness to make loans?"

Observed Frequencies

	Satisfied	Somewhat Satisfied	Dissatisfied
Idaho	23 (70%)	6 (18%)	4 (12%)
Montana	32 (78%)	5 (12%)	4 (10%)

X^2 (obtained) = 0.70
X^2 (critical) = 6.00
df= 2

Hypothesis #4: Perception of Bank's Knowledge of Economy

There was literally no association between the respondents' perception of their bank's knowledge of the local economy and banking type. Ninety-four percent of respondents in Idaho and 93 percent in Montana felt that their local bank was very knowledgeable or at least knowledgeable of the local economy (see Table 5).

TABLE 5. PERCEPTION OF BANK'S KNOWLEDGE OF LOCAL ECONOMY

"How knowledgeable is your bank about the local economy?"

Observed Frequencies

	Very Knowledgeable	Knowledgeable	Very Unknowledgeable
Idaho	22 (61%)	12 (33%)	2 (06%)
Montana	29 (63%)	14 (30%)	3 (07%)

X2 (obtained) = 0.09

X2 (critical) = 6.00
df= 2

CONCLUSIONS

Based on our findings and the reviewed literature, there is evidence that on balance branch banks in rural areas probably increase competition and local credit. Yet, there are some doubts about branch banking in rural areas since decisions about loans are often not made locally.

Our findings, based on surveys of small businesses in Idaho (a branch banking state) and Montana (primarily a unit banking state in 1993), indicate that branch banks in Idaho charge lower interest rates on loans and pay higher interest rates on money market accounts. The lower interest rates charged on loans are consistent with our first hypothesis. The higher interest rate paid on money market accounts was consistent with our second hypothesis although we found no difference in the rates paid on checking accounts, saving accounts, and certificates of deposit. There was no association between satisfaction with the bank's willingness to make loans and the type of bank. This is consistent with our third hypothesis. Finally, contrary to our fourth hypothesis, there was no association between the respondent's perception of their bank's knowledge of the local economy and the type of bank.

The data presented and the literature reviewed here suggest that nationwide branch banking will not necessarily have detrimental effects in non-metropolitan areas. Furthermore, it is likely that small independent banks will be able to compete in this changing environment. Rose (1986, p. 35) states, "the survivability of small, well managed, independent banks would not be seriously threatened by coexistence with nationwide banking organizations." There will remain opportunities for independent banks to fill niches in areas where service is lacking.

ENDNOTES

1. It is still up to states to decide whether to take part in interstate banking under this law. The vast majority of states will allow interstate banking. However, a handful of states, such as Texas, where populist sentiment has a distrust of big banks, may decide to opt out of interstate banking.
2. Geographical expansion raises similar concerns, whether carried out by BHC acquisitions or by branching. According to Lawrence and Klugman (1990: p. 1090), "Geographic market expansion in the banking industry typically takes place, with varying degrees of legal restriction, via merger, branching, instate, and now out-state holding company acquisition. The political arguments against the traditional modes of expansion are not substantively different from the current arguments against interstate banking."
3. For a comprehensive review of this literature, see Robert J. Tokle and Mark K. McBeth, "The Effect of Branch Banking on Rural Areas: A Survey." *Pennsylvania Economic Review,* vol. 5, number 2, 1996: 1-9.
4. Note that loan-to-deposit ratios are often used to measure the extent that banks lend out their deposits.
5. For a copy of the survey results, please contact Mark. K. McBeth, Department of Political Science, Idaho State University, Pocatello, Idaho 83209.

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IS THE BASEBALL-BETTING MARKET EFFICIENT? SOME RECENT EVIDENCE

Ravija Badarinathi* and Ladd Kochman**

Major League Baseball (MLB) is fertile ground for testing the efficiency of people's average economic judgments. The 1994 season-ending strike, for example, was based on the argument that players' salaries were out-of-line with clubs, revenues. A less publicized application of the *efficient market hypothesis* is the question of whether baseball's betting odds accurately reflect the teams' prospects of winning or losing.

Woodland and Woodland (1994), hereafter W&W, found that the Las Vegas odds for baseball tend to underestimate the chances that MLB underdogs have to beat their favored opponents. More specifically, the probabilities implied by the betting lines that underdogs will win their games were consistently lower than the actual probabilities derived from year-ending results. This preference for favorites caused W&W to conclude that the documented *favorite-longshot bias* in racetrack betting—namely, that longshots are overbet—was reversed in the baseball-betting market.

In addition to identifying the bias reversal, W&W devised a pair of opportunistic betting rules to test market efficiency. One strategy called for betting on all MLB underdogs during the period which revealed the bias against underdogs (1979-89) while the second proposed betting on only those 1979-89 underdogs in games where the line was no greater than -165/155. (Odds of -165/155 convey that $165 must be bet on the favorite to win $100 and that a $100 wager on the underdog returns $155.) The reasoning behind W&W's more selective rule was that the bias against underdogs was more pronounced among those teams which were only slight or moderate underdogs than among those which were heavy underdogs.

Despite the fact that observed probabilities for MLB underdogs exceeded their market-estimated counterparts (15 out of the 26 categories of lines between -110/100 and -300/260 and 10 out of the 12 categories between -110/100 and 165/155) and led to Z-values (2.02 and 2.61, respectively) significant at $p < .05$, W&W reported that neither strategy was profitable after commissions. While W&W might be criticized for comparing actual and estimated probabilities only at the end of the season and thereby ignoring interim events such as streaks and personnel changes, they could counter that baseball odds adjust to any meaningful developments and make measurement intervals unnecessary.

What does emerge as a methodological flaw is W&W 's decision to equate underdogs with longshots. It is not uncommon for baseball's best clubs (e.g., Atlanta and Cleveland) to be the underdog when playing another capable team on the road and facing a strong pitcher. The folly of designating the Braves or Indian as *longshots* might be amusing if it didn't act to undermine W&W's premise regarding the favorite-longshot bias. Careful scrutiny of W&W's results reveals that among heavy underdogs (i.e., lines from -170/160 to -300/260) expected probabilities were **greater** than actual probabilities in nine of the 14 categories. To the extent that only heavy underdogs should be regarded as longshots we would argue that W&W actually found support for the favorite-longshot bias and not the reversal they claimed.

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METHODOLOGY

The purpose of this study is threefold: (1) to test the W&W hypothesis that the baseball-betting market undervalues underdogs, (2) to place appropriate (and imaginary) bets on underdogs if the W&W hypothesis is confirmed and (3) to look for evidence of the favorite-longshot bias. Without data for the current or preceding season and eliminating the strike-shortened 1994 year, we chose the 1993 season as our test period. The source of betting lines and final scores was the *1993 Baseball Insight Log Book*—a publication available to the public through the Gambler's Book Club in Las Vegas.

Our initial step was to drop from consideration the games in which there was no underdog (i.e., *even* games) and to assign the remaining contests (out of the 2268 scheduled) to one of 30 different categories of lines ranging from -110/100 to -370/320. Subjective (or estimated) probabilities (SP) that underdogs would win were then computed for each of the 30 lines by using the W&W equation:

$$SP = 1/(b + 1) \tag{1}$$

where *b* is the average of the favorite's cost of winning \$1 (b_f) and the underdog's payoff from betting \$1 (b_u)—or $(b_f + b_u)/2$. For a team installed as a -165/155 underdog, *b* would be equal to 1.60—or (1.65 + 1.55)/2—and *SP* would, in turn, be equal to 0.3846—or 1/(1.60 + 1).

Underdogs' objective (or actual) probabilities (OP) of winning were next determined for the various line categories by summing the victories achieved by underdogs in a particular category at the end of the season and dividing the sum by the number of games assigned that line. Thus, if 100 games were handicapped -165/155 in 1993 and underdogs won 40 of them, the objective probability for that category would be 40 percent. (The terms *objective* and *subjective* introduced above were used by W&W and repeated here for the sake of continuity.)

By comparing objective and subjective probabilities for each category, it becomes possible to learn whether the success of underdogs surpasses market expectations—that is, whether OP > SP for a majority of cases—or vice versa. Of course, in an efficient market, objective and subjective probabilities for the same categories should be roughly equal. Statistical testing not unlike that used by W&W will disclose whether the two kinds of probabilities in 1993 differed to any exploitable degree. Furnishing the data for that testing will be the procedure below.

Step #1: Actual returns (R_a) for all line categories are calculated per the W&W equation:

$$R_a = OP\,(b_u + 1) - 1 \tag{2}$$

Step #2: An actual mean return ($\overline{R}_a$) is computed by multiplying the returns obtained from the preceding step by the categories' respective weights—or percent of total lines—and summing the resulting products.

Step #3: Expected returns (R_e) for all categories are derived by replacing *OP* in Eq. (2) with *SP*. Hence,

$$R_e = SP(b_u + 1) - 1 \tag{3}$$

Since the expected return in an efficient market is simply the bookmaker's commission, expected returns can alternately be found by:

$$R_e = (b_f - b_u)/(b_f + b_u + 2) \tag{4}$$

Step #4: An expected mean return ($\overline{R}_e$) is computed by multiplying the returns from Step #3 by the same weights employed in Step #2 and summing the products.

Step #5: Variances (VAR_L) for the expected returns connected with specific lines and derived in Step #3 are found by:

(5) $$VAR_L = (b_u + 1)^2 SP\,(l - SP)$$

Step #6: An overall variance (VAR_e) is calculated by multiplying the line variances from Step #5 by the weights used in Step #2, summing the products and dividing by the total number of lines.

As asserted earlier, an efficient baseball-betting market should not allow objective and subjective probabilities to be measurably different—nor the mean returns ($\overline{R}_a$ and $\overline{R}_e$, respectively) based on those probabilities. To that extent, we would expect the standardized test statistic (Z)

(6) $$Z = \frac{\overline{R}_a - \overline{R}_e}{SD_e}$$

(where SD_e is the square root of VAR_e from Step #6) not to be associated with probabilities less than five percent.

RESULTS

As revealed by Table 1, replication of W&W's rule to place wagers on all MLB underdogs produced a positive return ($\overline{R}_a$) of 1.69 percent vis-à-vis the expected return ($\overline{R}_e$) of -2.19 percent. When the variances for the 30 categories of lines in Table 1 were computed per Eq. (3), multiplied by respective weights, summed and finally divided by the 2165 total lines, an overall variance (VAR_e) of 6.244 percent emerged. Thus, after converting the rule's variance to a standard deviation of 2.50 percent and applying Eq. (6), it became apparent that betting on all MLB underdogs in 1993 resulted in a Z-value (1.55) that is significant at the 10-percent level.

(7) $$Z = \frac{.0169 - (-.0219)}{.0250} = 1.55$$

It is also evident from Table 1 that actual returns for the 30 categories of lines were more likely to be positive than negative when betting odds were closer to the smallest value (-110) than to the largest (-370). W&W had observed the same tendency during the 1979-89 years and attempted to exploit it by limiting their imaginary bets to underdogs in games where the lines were -165/155 or lower. When we replicated W&W's more selective strategy over the 1993 season, our Z-value declined. Although objective probabilities for the 12 lines ranging from -110 through -165 exceeded their subjective counterparts in nine cases, the actual return ($\overline{R}_a$) slipped to 1.11 percent (from the 1.69 percent earned by the first rule) while the expected return (Re) was unchanged at -2.19 percent. When the variances for those 12 categories of lines in Table 2 were computed per Eq. (5), multiplied by respective weights, summed and then divided by the 1728 total lines, a slightly higher variance (7.13 percent) and standard deviation (2.67 percent) surfaced and a lower and insignificant Z-value (1.24) resulted.

(8) $$Z = \frac{.111 - (-.0219)}{.0267} = 1.24$$

TABLE 1. ACTUAL ($\bar{R}_a$) AND EXPECTED MEAN RETURNS ($\bar{R}_e$) FROM BETTING ON ALL MLB UNDERDOGS IN 1993

LINE	(2) NO. OF LINES	(3) LINES AS % OF TOTAL	(4) OBJECTIVE PROB.	(5) SUBJECTIVE PROB.	(6) ACTUAL RETURN	(7) COLUMNS (3) X (6)	(8) EXPECTED RETURN	(9) COLUMNS (3) X (8)
-110	190	.0878	.5474	.4878	.0948	.0086	-.0244	-.0021
-115	168	.0776	.5000	.4762	.0250	.0019	-.0238	-.0018
-120	180	.0831	.4500	.4651	-.0550	-.0046	-.0233	-.0019
-125	150	.0693	.5200	.4545	.1180	.0082	-.0227	-.0016
-130	164	.0758	.3780	.4444	-.1683	-.0127	-.0222	-.0017
-135	162	.0748	.4506	.4348	.0139	.0010	-.0217	-.0016
-140	157	.0725	.4586	.4255	.0548	.0040	-.0213	-.0015
-145	143	.0661	.4336	.4167	.0190	.0013	-.0208	-.0014
-150	131	.0605	.3969	.4082	-.0474	-.0029	-.0204	-.0012
-155	83	.0383	.4337	.4000	.0626	.0024	-.0200	-.0008
-160	114	.0527	.3947	.3922	-.0133	-.0007	-.0196	-.0010
-165	86	.0397	.4186	.3846	.0674	.0027	-.0192	-.0008
-170	79	.0365	.3671	.3774	-.0455	-.0017	-.0189	-.0007
-175	41	.0189	.4878	.3704	.2927	.0055	-.0185	-.0003
-180	65	.0300	.4000	.3636	.0800	.0024	-.0182	-.0005
-185	36	.0166	.5000	.3571	.3750	.0062	-.0179	-.0003
-190	40	.0185	.4000	.3509	.1200	.0022	-.0175	-.0003
-200	50	.0231	.3600	.3419	.0260	.0006	-.0256	-.0006
-210	35	.0162	.2571	.3333	-.2544	-.0041	-.0333	-.0005
-220	30	.0139	.2667	.3226	-.1999	-.0028	-.0323	-.0004
-230	12	.0055	.4167	.3125	.2917	.0016	-.0313	-.0002
-240	8	.0037	.3750	.3030	.2000	.0007	-.0303	-.0001
-250	9	.0042	.2222	.2941	-.2667	-.0011	-.0294	-.0001
-260	15	.0069	.2667	.2857	-.0932	-.0006	-.0286	-.0002
-270	3	.0014	.3333	.2778	.1666	.0002	-.0278	.0000
-280	4	.0018	.2500	.2703	-.1000	-.0002	-.0270	-.0001
-300	1	.0005	.0000	.2632	-1.0000	-.0005	-.0526	.0000
-320	4	.0018	.5000	.2500	.9000	.0016	-.0500	-.0001
-350	4	.0018	.0000	.2353	-1.0000	-.0018	-.0588	-.0001
-370	1	.0005	.0000	.2247	-1.0000	-.0005	-.0563	.0000
	2165					$\bar{R}_a$=.0169		$\bar{R}_e$=-.0219

TABLE 2. ACTUAL ($\bar{R}_a$) AND EXPECTED MEAN RETURNS ($\bar{R}_e$) FROM BETTING ON 1993 MLB UNDERDOGS IN GAMES WITH LINES </= -165

LINE	(2) NO. OF LINES	(3) LINES AS % OF TOTAL	(4) OBJECTIVE PROB.	(5) SUBJECTIVE PROB.	(6) ACTUAL RETURN	(7) COLUMNS (3) X (6)	(8) EXPECTED RETURN	(9) COLUMNS (3) X (8)
-110	190	.1100	.5474	.4878	.0948	.0104	- .0244	- .0027
-115	168	.0972	.5000	.4762	.0250	.0024	- .0238	- .0023
-120	180	.1042	.4500	.4651	- .0550	- .0057	- .0233	- .0024
-125	150	.0868	.5200	.4545	.1180	.0102	- .0227	- .0020
-130	164	.0949	.3780	.4444	- .1683	- .0160	- .0222	- .0021
-135	162	.0938	.4506	.4348	.0139	.0013	- .0217	- .0020
-140	157	.0909	.4586	.4255	.0548	.0050	- .0213	- .0019
-145	143	.0828	.4336	.4167	.0190	.0016	- .0208	- .0017
-150	131	.0758	.3969	.4082	- .0474	- .0036	- .0204	- .0015
-155	83	.0480	.4337	.4000	.0626	.0030	- .0200	- .0010
-160	114	.0660	.3947	.3922	- .0133	- .0009	- .0196	- .0013
-165	86	.0498	.4186	.3846	.0674	.0034	- .0192	- .0010
	1728					$\bar{R}_a$ = .0111		$\bar{R}_e$ = -.0219

CONCLUSIONS

It seems fair to conclude that bettors continue to underestimate the chances of baseball's underdogs. Objective probabilities for underdogs beat subjective probabilities in 60 percent of the 30 line categories. Among slight-to-moderate underdogs, the advantage increased to 75 percent of the 12 corresponding categories.

While the aforementioned imbalance of probabilities generated positive returns for a majority of betting lines, no significant Z-values at $p < .05$ emerged. It appears that the baseball-betting market is able to absorb certain patterns of behavior on the part of its participants and remain another example of the *efficient market hypothesis.*

Finally, while the results of this study parallel that performed by W&W, we make no assertion of a reversed favorite-longshot bias. In fact, by defining longshots as underdogs in games with odds of -170/160 or longer and finding that *OP > SP* in exactly half of the 18 affected categories, we are inclined to argue that the favorite-longshot bias simply does not exist among baseball bettors.

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THE EFFECT OF OPTIONS LISTING STATUS ON TRADING VOLUME REACTION TO EARNINGS ANNOUNCEMENTS

Edward J. Conrad* and R. Penny Marquette*

INTRODUCTION

Prior studies, such as those by Beaver (1968), Kiger (1972), and Morse (1981), document an increase in trading volume around earnings announcements. Variables that may be linked with this elevated trading volume, such as firm size (Zeghal, 1984 and Bamber, 1986 and 1987) and the dispersion of financial analysts' earnings forecasts (Ziebart, 1990; Atiase and Bamber, 1994; and Bamber and Cheon, 1995) have been the subject of empirical scrutiny. This study will examine the effect of options listing status on the trading volume reaction to earnings announcements. Unfortunately, there is no comprehensive theory of trading volume that can serve as a guide in the empirical investigation. Ross (1989, p. 94) states that, "There is no serious research yet on volume data. This is in large part because we have no serious theories that purport to explain the volume of trade...." It is hoped that the results of this study will provide grist for the theorists' mill.

HYPOTHESIS

This paper explores the hypothesis that trading volume in the common stock of an optionable firm will be different from that of a non-optionable firm. We will employ a non-directional test of this hypothesis since there is little theoretical or empirical evidence to support this assumption.

To compare optionable firms and non-optionable firms it is necessary to match them on several dimensions. A count of articles appearing in *The Wall Street Journal* shows that larger firms receive greater news coverage. We therefore group firms by size in order to mitigate the effects of the predisclosure of information. They must also be matched in time, and we accomplish this by examining the period surrounding earnings announcement dates. Equally important is the nature of the earnings announcement. Prior research (Lakonishok and Smidt, 1986, and Jain and Joh, 1988) indicates that earnings announcements construed by the market as bad news are associated with lower trading volume around the announcement date (for the same absolute amount of unexpected earnings) when compared with good news announcements. A possible explanation for the difference in market reaction between good news and bad news is that short sale restrictions limit the ability of investors to react to negative information. Karpoff (1988) states that short sales are relatively costly for at least two reasons: a short sale can only be made on an uptick, or a zero-plus tick (this is true for exchange listed stocks, but not for Over the Counter (OTC) stocks); and the short sale proceeds are not remitted to the short seller, causing interest to be foregone on the proceeds, and yet the short seller remains liable for any dividends paid on the stock.

Options trading in a firm's stock may result in considerations that are related to such short sale restrictions. Perhaps options trading provides an outlet for bad news trading that would otherwise be

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delayed (or foregone). Diamond and Verrecchia (1987) state that the introduction of options trading reduces the cost of short selling, resulting in increased speed of adjustment to private information, especially to bad news. Based on the foregoing discussion, it would appear necessary to control for the good news/bad news nature of the earnings announcement when examining trading volume differences between option and nonoption firms. Accordingly, we will classify firms according to the quality of the earnings announcements into "good news" and "bad news" groups.

Insight into the relative roles played by the availability of options trading, the quality of earnings announcements, and environmental and firm-specific characteristics will enrich our understanding of the role played by both options and new information in the capital markets.

PREVIOUS STUDIES

Differential security price reaction to earnings announcements based upon option versus nonoption listing status has been examined in prior work. Jennings and Starks (1986) find quicker intraday adjustment to earnings announcements for option than for nonoption firms. Skinner (1990) compares stock price adjustments to earnings announcements before and after options trading is introduced and finds smaller intrafirm price adjustment to quarterly earnings announcements after the introduction of options trading in the common shares. He suggests that the introduction of options trading increases the level of investor interest and the amount of predisclosure information. Ho (1993) finds greater abnormal return volatility around earnings announcements for nonoption firms than for option firms.

There is little direct empirical evidence pertaining to the relationship between the trading volume in a firm's stock and the availability of options trading in its stock. Anthony (1988) compared the speed of reaction of stock volume and option volume to earnings announcements and concluded that the trading in options leads the trading in the underlying common shares by one day. In a follow-up study, Stephan and Whaley (1990) note since options markets close ten minutes after the stock markets, this may explain Anthony's findings. Correcting for the difference in closing times, Stephan and Whaley conclude that stock market trading actually leads trading in the options market.

Two studies comparing the behavior of a firm's stock before and after option trading begins are Whiteside, Dukes, and Dunne (WDD) (1983) and Hayes and Tennenbaum (1979). Both studies consider the differences in stock volume caused by the existence of options trading. Unfortunately, they arrive at conflicting results. Hayes and Tennenbaum (1979) conclude that options trading does not result in increased trading volume for the firm's stock, while WDD (1983) find that the presence of options trading does increase the trading volume of the underlying common shares.

RESEARCH DESIGN

<u>Data Sources and Sample Selection</u>

The Media General Price-Volume (MGPV) Data Tape supplied firm-specific trading volume data. The MGPV includes only those firms that have a complete data set for the period July 11, 1976 through February 13, 1987.[1] The number of shares outstanding (for calculating percentage trading volumes) was obtained from the Center for Research in Security Prices (CRSP) data tape; firm size (calculated as the beginning fair market value of common shares outstanding) and earnings announcement dates were taken from the Standard and Poor's Annual Compustat data base. New York Stock Exchange (NYSE), American Exchange (AMEX), and OTC firms with annual earnings announcements were matched for data availability with the above data sources during the six year period from 1980 through 1985. This time period was selected due to data availability constraints.[2]

All firms which had the necessary Compustat, MGPV, and CRSP data for the interval chosen were selected for the initial sample. Each observation was then examined, using the *Wall Street Journal*, to ensure that there were no confounding events in the five days before and after each earnings announcement (the "event window"). Confounding events included news involving mergers and acquisitions, strikes or settlements, debt restructurings or ratings changes, and significant litigation. There were 1,254 individual firms remaining in the final sample. Because multiple earnings announcements were examined for each firm, the final sample consisted of 4,300 observations.

The availability of options trading in a firm's common shares was determined by examining options activity listed in the *Wall Street Journal* for an arbitrarily selected day in January of each year. If a firm had a listed option at the beginning of the year in which it made its annual earnings announcement, as well as at the beginning of the next year, it was designated as a firm for which options trading existed. If the firm had no such option listing at these two times it was designated as a firm for which options trading did not exist.[3]

Descriptive Statistics

Excessive industry concentration does not appear to be a problem for this sample. The 1,254 firms belong to 295 different four digit SIC code industry groupings, with an average of 4.25 firms in each SIC code, and a median of 3.0.

Table 1 provides a description of the 4,300 observations in the sample by year and market. Approximately 77 percent of the 4,300 observations are NYSE firms, approximately 13 percent are AMEX firms, and the balance are OTC firms. The firms are mainly calendar year firms with fiscal years ending December 31. The balance, approximately 32 percent, have non-calendar year ends.

TABLE 1. DISTRIBUTION OF EARNINGS ANNOUNCEMENTS BY MARKET AND YEAR[a]

	1980	1981	1982	1983	1984	1985	Total
NYSE:	460	541	567	578	576	583	3,305
AMEX:	58	81	95	100	111	110	555
OTC:	49	56	63	89	91	92	440
Total	**567**	**678**	**725**	**767**	**778**	**785**	**4,300**

[a] New York Stock Exchange (NYSE), American Stock Exchange (AMEX), and Over the Counter (OTC) firms.

Classification as Good News or Bad News Earnings Announcements

The relevant portion of the earnings announcement for explaining both abnormal price and abnormal volume reactions, is the *unexpected* component. There is considerable theoretical basis for estimating *anticipated* earnings or volume, comparing that estimate to actual returns and volume, and creating the test metrics of Cumulative Abnormal Return (CAR) and Cumulative Abnormal Volume (CAV). Background can be found in work by Foster, Olsen, and Shevlin (1984), Brown, Foster, and Noreen (1985), Bhushan (1989), and Ajinkya and Jain (1989).

As is typically done, this study will estimate the security market's expected annual return surrounding the earnings announcement date using the market model. The *unexpected* portion of the security return metric will be the residual from a market model regression of the form:

(1) $$r_{it} = \alpha_i + \beta_i(R_{mt}) + \varepsilon_{it},$$

where:

r_{it} = return for firm i in period t,

α_i = intercept for firm i,

β_i = slope coefficient for firm i,

R_{mt} = equally-weighted market returns for all NYSE and AMEX firms, with a separate market return calculated for the OTC firms,

ε_i = error term satisfying the assumptions of the standard linear regression model.

The return model will be estimated over a 238 day estimation period (119 days on either side of the event window). The same time interval will be used to estimate the market model for trading volume. The event window itself will encompass the five days on either side of the earnings announcement, for a total of eleven days. The CAR will be calculated as:

(2) $$CAR_i = \prod_{t}^{T}(1+\varepsilon_{it}) - 1.$$

When the actual return exceeds the anticipated return over the period of interest and the CAR is positive, the security will be placed in a "good news" category for that earnings announcement. Conversely when the anticipated return exceeds the actual return and the CAR is negative, the earnings announcement will be placed in a "bad news" category.

Trading Volume Metric

This study makes use of the results of Ajinkya and Jain (1989) in specifying the normal trading volume metric.[4] The procedure for creating a variable for abnormal trading volume is similar to that for determining abnormal security returns. "Normal" trading volume is estimated for the event window, compared with actual trading volume, and the difference summed into a Cumulative Abnormal Volume (CAV).

The volume trading index used to estimate normal trading volume will consist of an equally-weighted market index designed to minimize the influence of large firms.[5] The equally-weighted market index specification is calculated as shown below. The *N* in equation (3) represents the number of firms across all markets.[6]

(3) $$EWP = \frac{\sum_{n=1}^{N} \frac{\text{Log(Number of Shares Traded} + 1)}{\text{Log(Number of Shares Outstanding} + 1)}}{N}.$$

The regression will be estimated over the 119 days period on either side of the eleven day event window centered on the annual earnings announcement. The total (238 day) estimation period will always exclude the eleven day event window centered on any quarterly earnings announcement that occurred within the model estimation period.[7]

The abnormal volume for a firm in the earnings announcement period will be a residual from a volume market model of the form:

$$\ln(V_{it}) = \alpha_i + \beta_i(V_{mt}) + \varepsilon_{it}, \tag{4}$$

where:

V_{it} = percentage of firm i's shares traded on day t,

α_{it} = intercept for firm i,

β_i = slope coefficient for firm i,

V_{mt} = percentage of market shares traded on day t per equation (3),

ε_{it} = volume residual for firm i on day t.

The CAV over any portion of the event window is simply the sum of the volume residuals for the days in question.

Table 2 provides the distribution of earnings announcements by firm size and the quality of the announcements.

TABLE 2. EARNINGS ANNOUNCEMENTS DISTRIBUTED BY FIRM SIZE, EARNINGS ANNOUNCEMENT QUALITY, AND OPTION STATUS[a]

	Optionable		Non-Optionable	
Size Decile	**Good News**	**Bad News**	**Good News**	**Bad News**
1	0	0	204	226
2	0	0	224	206
3	0	0	230	200
4	5	4	170	251
5	10	14	187	219
6	21	22	207	180
7	46	38	146	200
8	61	74	149	146
9	103	99	105	123
10	161	187	45	37
Totals:	**407**	**438**	**1,667**	**1,788**

[a] Firm size calculated as beginning of fiscal year fair market value of common stock outstanding. News classification based on three day (-1,0,+1) cumulative abnormal return from equally-weighted market model.

RESULTS

Data

Tables 3 and 4 present the mean Cumulative Abnormal Return (CAR) and Cumulative Abnormal Trading Volume (CAV) for all firms, by size decile, during the three days centered on the earnings announcement. Overall, there is not a statistically significant difference in the magnitude of absolute price reaction for option firms or nonoption firms between good news and bad news. There is also not a statistically significant difference in the magnitude of absolute price reaction for good news and bad news between option firms and nonoption firms. These results are not directly comparable to those of prior studies because prior studies did not control for the good news/bad news nature of the earnings announcement.

TABLE 3
MEAN CUMULATIVE ABNORMAL RETURN (CAR)[a] BY SIZE, GOOD NEWS (GN)/BAD NEWS (BN), AND OPTION STATUS

Size Decile	Overall		Optionable		Non-Optionable	
	GN	BN	GN	BN	GN	BN
1. CAR	.0522	-.0429	***	***	.0522	-.0429
2. CAR	.0407	-.0369	***	***	.0407	-.0369
3. CAR	.0304	-.0363	***	***	.0304	-.0363
4. CAR	.0335	-.0275	.0279	-.0573	.0036	-.0271
5. CAR	.0305	-.0258	.0465	-.0369	.0296	-.0251
6. CAR	.0286	-.0252	.0297	-.0312	.0285	-.0245
7. CAR	.0259	-.0290	.0321	-.0405	.0240	-.0269
8. CAR	.0314	-.0258	.0384	-.0245	.0286	-.0265
9. CAR	.0262	-.0272	.0288	-.0309	.0237	-.0242
10. CAR	.0232	-.0269	.0228	-.0272	.0246	-.0253

[a]Three day (-1,0,+1) CAR determined by equally-weighted market model.
***No optionable firms.

TABLE 4. MEAN CUMULATIVE ABNORMAL VOLUME (CAV)[a] BY SIZE, GOOD NEWS (GN)/BAD NEWS (BN), AND OPTION STATUS

Size Decile	Overall		Optionable		Non-Optionable	
	GN	BN	GN	BN	GN	BN
1. CAV	.1363	.1197	***	***	.1363	.1197
2. CAV	.1051	.0725	***	***	.1051	.0725
3. CAV	.0964	.0605	***	***	.0964	.0605
4. CAV	.1097	.0429	.0054	-.0180	.1127	.0439
5. CAV	.0793	.0422	.1828	.0109	.0737	.0442
6. CAV	.0543	.0194	.0559	.0115	.0542	.0204
7. CAV	.0315	.0333	.0233	.0214	.0340	.0356
8. CAV	.0477	.0393	.0758	.0459	.0362	.0359
9. CAV	.0390	.0269	.0560	.0393	.0223	.0168
10. CAV	.0318	.0259	.0409	.0365	-.0010	-.0274

[a]Three day (-1,0,+1) CAV determined by overall equally-weighted market model.
***No optionable firms.

The pattern for trading volume (Table 4) yields interesting results. Within options trading status, "good news" abnormal trading volume exceeds that for "bad news" in all size deciles with the exception of size decile 7 for non-optionable firms. Comparison of "good news" abnormal trading volume between options listing status categories reveals that size decile 7 is again an exception to the apparent pattern. Ignoring size decile 4, which contains only 9 option firms versus 421 nonoption firms, the trading volume for "good news" announcements is higher for option firms than for nonoption firms in all size deciles except size decile 7. In the "bad news" group, we find mixed results with the fifth, sixth and seventh deciles showing larger trading volume for the non-optionable firms, while optionable firms exhibit larger CAVs in deciles 8, 9 and 10.

Further examination of Table 2 indicates an almost perfect monotonic relationship between firm size and the likelihood that options trading in its common shares will be available. As we have seen, there are no option firms in the smallest three size deciles and the vast majority of option firms are concentrated in the three largest size deciles. These three groups yield an approximately equal number of option (685) and nonoption (605) firms which will be used in the analyses that follow.[8]

Prior research (see, e.g., Black, 1975 and Diamond and Verrecchia, 1987) suggests that options increase the amount of information available to the market due to investors' increased incentives to collect private information. If true, we might expect to see significantly larger trading volume in the very early (pre-announcement) event period for firms with options. This result may be more pronounced for "good news" than for "bad news" firms for the reasons cited earlier, but optionable firms should show a clear lead in the reaction to new information. We can examine this conjecture by looking at the very early event window—the days preceding the earnings announcement.

Evidence of improved liquidity for optionable firms may also be found in the late event period, those days immediately after the earnings announcement. It is here that the division of the earnings announcements into "good news" and "bad news" becomes particularly helpful. For "good news" firms we would expect those *with* options to demonstrate a lower trading volume in the late event period (assuming that the optionable firms have already responded to pre-announcement information). Similarly, for optionable "bad news" firms, we might also expect less volume if, in fact, option trading was used as a substitute for short sale activity in the pre-announcement period.

The Figure below defines the questions to be addressed in the analysis that follows. To answer questions about the impact of options trading in the market, it is necessary to examine the relationships both across the columns and across the rows.

Optionable – Good News	Non-Optionable – Good News
Optionable – Bad News	Non-Optionable – Bad News

It is obvious that we want to know whether optionable firms and non-optionable firms react differently (across the columns), but it is also interesting to know whether the good news/bad news dichotomy affects optionable firms differently from non-optionable firms (across the rows). Finally, we will examine three event windows: early (t = -1 to +1), late (t = +1 to +3) and a finally, a very early window (t = -2 to -3), for insight into the effect of options trading on market reaction times.

Tables 5 and 6 present the comparison of cumulative abnormal volume in the early (Table 5) and late (Table 6) event windows. In the early event window (t = -1 to +1) the CAV reaction to "good news" is greater than that for "bad news" within option listing status, but the difference is not statistically significant. The reaction to the quality of the earnings announcements has either taken place earlier, or is delayed until a later period. We do, however, find a weakly significant difference between the optionable and non-optionable firms, with optionable firms showing a greater reaction to *both* "good" and "bad" news. Table 6

TABLE 5. CUMULATIVE ABNORMAL VOLUME (CAV) ANALYSIS
EARLY EVENT WINDOW (t = −1 TO +1)
(FIRMS FROM SIZE DECILES 8, 9, AND 10)

	Optionable		Non-Optionable	
	Good News (n=325)	Bad News (n=360)	Good News (n=299)	Bad News (n=306)
Mean CAV	.0530	.0393	.0257	.0204
Standard Deviation	(.1095)	(.1158)	(.1483)	(.1572)

Wilcoxon Tests of Significance:	*p* value[a]
Option versus Nonoption Firms reporting Good News:	.007
Option versus Nonoption Firms reporting Bad News:	.123
Option firms only, Good News versus Bad News:	.103
Nonoption firms only, Good News versus Bad News:	.814

[a]The *p* values report the probability that the means of the two groups are equal, based on a two-tailed, two sample, Wilcoxon means test. Thus, the probability that options firms react to good news in the same manner as nonoption firms is .007.

TABLE 6. CUMULATIVE ABNORMAL VOLUME (CAV) ANALYSIS
LATE EVENT WINDOW (t = +1 TO +1)
(FIRMS FROM SIZE DECILES 8, 9, AND 10)

	Optionable		Non-Optionable	
	Good News (n=325)	Bad News (n=360)	Good News (n=299)	Bad News (n=306)
Mean CAV	.0356	.0143	.0231	.0081
Standard Deviation	(.1082)	(.1058)	(.1610)	(.1411)

Wilcoxon Tests of Significance:	*p* value[a]
Option versus Nonoption Firms reporting Good News:	.189
Option versus Nonoption Firms reporting Bad News:	.302
Option firms only, Good News versus Bad News:	.014
Nonoption firms only, Good News versus Bad News:	.115

[a]The *p* values report the probability that the means of the two groups are equal, based on a two-tailed, two sample, Wilcoxon means test.

shows that the CAV is again greater for option than for nonoption firms within the late event window (t = +1 to +3) for both "good" and "bad" news, but the difference is not statistically significant. However, the CAV reaction to "good news" is significantly greater than that for "bad news" within option listing status. It is not the case, as previously conjectured, that option firms exhibit lower abnormal trading volume than nonoption firms in the late event window.

Table 7 steps back to an earlier time period in the event window (t = -2 to -3). Here we find evidence that, within option listing status, "bad news" earnings announcements are associated with significantly greater abnormal volume behavior than are "good news" earnings announcements. Additionally, CAV is significantly greater for option than for nonoption firms within this event window for both "good" and "bad" news. As hypothesized by Diamond and Verrecchia (1987), the existence of options trading is associated with an increased speed of adjustment (as measured by abnormal volume) to private information, especially to "bad news."

TABLE 7. CUMULATIVE ABNORMAL VOLUME (CAV) ANALYSIS MARKET REACTIONS PRIOR TO EARNINGS ANNOUNCEMENTS (t = -2 TO -3) (FIRMS FROM SIZE DECILES 8, 9, AND 10)

	Optionable		Non-Optionable	
	Good News (n=325)	Bad News (n=360)	Good News (n=299)	Bad News (n=306)
Mean CAV	.0010	.0115	–.0215	–.0005
Standard Deviation	(.0743)	(.0846)	(.1133)	(.0990)

Wilcoxon Tests of Significance:	*p* value[a]
Option versus Nonoption Firms reporting Good News:	.010
Option versus Nonoption Firms reporting Bad News:	.062
Option firms only, Good News versus Bad News:	.100
Nonoption firms only, Good News versus Bad News:	.043

[a]The *p* values report the probability that the means of the two groups are equal, based on a two-tailed, two sample, Wilcoxon means test.

The hypothesis of Diamond and Verrecchia is limited to *option effects* on the speed of market adjustment, but we find that the phenomenon of increased market reaction prior to "bad news" earnings announcements is not limited to the consideration of options status. No matter how the data are selected—OTC versus exchange, option versus nonoption, or by size—we find that market reactions are more pronounced for "bad news" in the pre-announcement window than for "good news." Although not presented in Table 7, this result was particularly prominent for OTC firms, the vast majority of which (430 of 440) are not optionable.[9] We consider this evidence that factors beyond the existence of options must be considered when comparing "good" and "bad" news market reactions.

A possible explanation for the larger abnormal volume reaction to "bad news" earnings announcements prior to the earnings announcement date, and the larger abnormal volume reaction to "good news" earnings announcements on (Table 5) and after (Table 6) the earnings announcement date, may be provided by timing studies such as those by Penman (1984), Chambers and Penman (1984), and Kross and Schroeder (1984). Firms tend to make "good news" earnings announcements more promptly than "bad news" earnings announcements. Therefore, the market reaction to "bad news" earnings announcements may begin sooner than for "good news" announcements because the market is anticipating the "bad news" announcement prior to its actual release. This preemption of the information content of "bad news" earnings announcements, along with the more timely nature of "good news" earnings releases, also explains the comparative differences in abnormal volume reactions on and after the actual earnings release date.

Table 8 further illustrates the point that factors beyond the existence of options must be considered when attempting to explain market reactions to earnings announcements. We know that the option firms are, on average, larger. But an examination of Table 8 indicates that they are also riskier (with a beta of 1.286 compared with 1.019), draw more analysts' attention, and are mentioned far more frequently in the business press (43.3 compared with 19.9 citations in the *Wall Street Journal).*

TABLE 8
COMPARISONS—OPTION AND NONOPTION FIRMS
(FIRMS FROM SIZE DECILES 8, 9, AND 10)

	Option	Nonoption
Fair Market Value of Common Stock Outstanding in Millions of Dollars[a]	$3,801.6	$1,505.6
(Standard Deviation)	(6,177.8)	(3,999.9)
Financial Analysts[b]	20.6	15.1
(Standard Deviation)	(7.3)	(5.9)
Wall Street Journal Articles[c]	43.3	19.9
(Standard Deviation)	(47.3)	(11.8)
Beta[d]	1.286	1.019
(Standard Deviation)	(.436)	(.527)

[a] Beginning of the period.

[b] Number of financial analysts comprising the Institutional Brokers' Estimate System (I/B/E/S) forecast for the nearest month preceding the month of the annual earnings announcement.

[c] Number of article citations in the *Wall Street Journal Index* since the last annual earnings announcement.

[d] Betas from market model regressions employing equally weighted indices.

CONCLUSIONS

There is clearly a difference between optionable and non-optionable firms. The optionable firms are larger, receive more analyst and press attention, and their share activity reacts more quickly to "bad news" in the pre-announcement period (t = -2 to -3), and with more volatility to "good news" in the early (t= -1 to +1) and late (t = +1 to +3) post-announcement periods.

Stock exchanges value the ability to offer options in the stocks of firms for which the underlying securities are volatile. That volatility would likely also lead to increased analyst and press attention as would larger size. The question that remains to be answered is whether option firms are different because they offer options, or whether they offer options because they are different.

ENDNOTES

1. Firms that went out of business or were acquired during this period are not included on the MGPV. Clearly, this imparts a survivorship bias to the results of this study.
2. Investigation leads us to believe that the data tape employed in this study may be the last (or very close to the last) produced by Media General.
3. Instances in which options trading was found at only one of the two bracketed times (154 cases) were resolved by reference to microfilm copies of the *Wall Street Journal* options listing page on the annual earnings announcement date.
4. Ajinkya and Jain (1989) provide a comprehensive analysis of the properties of raw trading volume and the residuals obtained from mean-adjusted and market model specifications of normal trading volume. They conclude that a log transformation of the raw trading volume is approximately normally distributed, and that a market model approach is superior to a mean-adjusted approach in specifying normal volume. First order autocorrelation is present in trading volume residuals, and they find that an estimated generalized least square (GLS) model with an AR(1) structure imposed on the residuals yields an improvement in market model specification.
5. This metric will be constructed as the average percentage of shares traded for all firms in the market index for a given day, and this average across all firms used as the independent variable in a regression in which the dependent variable is the percentage of a firm's outstanding shares traded on a given day.
6. A single share is added to the number of shares traded and outstanding for each firm to avoid the problem of taking the logarithm of zero. Calculation of the market index will be based on a GLS (regression) approach that will take into account the first order serial correlation in the volume market model regression residuals.
7. Although the period employed in model estimation excludes the eleven day period surrounding earnings announcements (assuring that any residuals from the model in earnings announcement periods will be more clearly defined as resulting from the earnings announcement event), it is necessary that these days be included in an estimation employing the GLS approach. For example, the preceding day's volume is necessary to implement the AR(a) structure specified in the GLS procedure. This is not a problem for nonevent days or for the first day in the event window. However, subsequent days in the event window will be contaminated by the event under investigation, making it desirable to use predicted values from the GLS procedure as the expected values on these days. Specifically, the earnings announcement days are included as missing observations in determining the GLS model parameters, and the GLS procedure computes the predicted value for these missing observations using the nonmissing observations.

   The predicted values allow the calculation of forecast errors for the annual earnings announcement event window. The GLS approach will have filled in the missing event window values with predicted values and the prediction errors can be calculated as the actual observed values less these predicted values. The CAV will simply be the total of these daily prediction errors over the event window of interest.
8. Including the fourth largest decile increases the number of option firms to 769, but heavily weights the sample toward nonoption firms (951). This clear size effect underlies the importance of controlling for firm size when conducting this or similar types of market research.
9. This may be explained by the fact that, as mentioned earlier, OTC firms are not subject to the same limitations on short selling as are exchange listed firms.

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CAUSALITY AND COINTEGRATION BETWEEN EXPORTS AND ECONOMIC GROWTH: A MULTIVARIATE APPROACH

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I. INTRODUCTION

The relationship between exports and economic growth has long been a subject of interest to economists as well as policy makers. Standard international trade theories suggest that exports can serve as an engine of growth by (i) reallocating resources according to comparative advantage, (ii) making it possible for both developed and developing countries to overcome the limitations of the domestic market and allowing them to take advantage of economies of scale, and (iii) enhancing productivity through technology transfer, innovations, and additional investment to take advantage of the enlarged market. Recently, Krueger (1980), Lucas (1990) and Ghartey (1993) have argued that export expansion generates increasing returns to scale and dynamic spill-over effects from the export sector to other sectors. Diffusion of state-of-the-art technology from the export sector to other sectors also takes place.

However, the role of exports in economic growth is not without controversy in the empirical literature. Cross-section studies by Emery (1967), Michaely (1977), Tyler (1981), Kavoussi (1984), Ram (1985), Moschos (1989), Cheng (1992), and others confirm the hypothesis that export expansion leads to faster economic growth. Although these regression studies make a significant contribution to the empirical literature, their shortcoming is that they attempt to equate correlation with causation. Another issue concerning regression specification is that these previous studies use cross-sectional data, which implicitly assume the existence of an invariant cross-section relationship between exports and economic growth. However, differences in the level of economic development across countries make this assumption questionable.

Time series studies (Jung and Marshall, 1985; Chow, 1987; Kunst and Marin, 1989; Bhamani-Oskooee, Modtadi, and Shabsigh; 1991; Ghartey, 1993; Bahmani-Oskooee and Alse, 1993, Dodaro, 1993; Dutt and Ghosh, 1996 and others), produce, however, an inconclusive outcome. While these time series studies have important merits, we find some deficiency in each approach:

(i) Non-stationarity problems (e.g., Jung and Marshall, 1985; Chow, 1987). A variable is stationary if its mean value and its variance do not change systematically over time (Gujarati, 1996). When estimating a model with nonstationary time series data, the usual ordinary least squares significance tests performed on the regression coefficients can be quite misleading (Engle and Granger, 1987). A more detailed discussion of this subject is given in Section II.

(ii) Ad hoc approach in selecting lags (e.g., Jung and Marshall, 1985; Chow, 1987). This refers to the practice of selecting the lag length arbitrarily when performing the causality test. To minimize biases in causality estimation, it is necessary to employ statistical methods such as final prediction error (FPE) because the specification does greatly affect the causality test results.

(iii) Entering regressors into the equation in arbitrary order (e.g., Ghartey, 1993). When using multivariate approaches, the causality results are affected by the sequence in which the regressors are entered into the equation (Caines, Keng, and Sethi, 1981).

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(iv) Omission of relevant variables (e.g., Bahmani-Oskooee and Alse, 1993). An omission of relevant variables may cause biases in causality estimation (Lutkepohl, 1982). This occurs when a bivariate model is employed because economic growth can be affected by many factors other than exports.

(v) The failure to use Johansen's test of multi-cointegration (e.g., Ghartey, 1993). In contrast to the Engle-Granger two-step cointegration test (Engle and Granger, 1987), Johansen's procedure (Johansen, 1991; Johansen and Juselius, 1990) should be used when a multivariate model is employed, as discussed in section II.

(vi) Failure to incorporate error-correction terms to recapture the lost long-run information in a differencing process (e.g., Jung and Marshall, 1985; Chow, 1987; Ghartey, 1993). The concept of cointegration examines the presence or absence of a long-run equilibrium relationship between two or more variables (Engle and Granger, 1987). Even though two or more variables are cointegrated, they may temporarily depart from the equilibrium state. With an error-correction model, a proportion of the disequilibrium in one period is corrected in the next period (Granger, 1988; Miller, 1992). More discussion of this subject is provided in Section II.

Any of these shortcomings can produce misleading and/or invalid causality inferences. As such, if the direction of causality between exports and economic growth is not clear, it may inhibit effective public policy-making. A need therefore exists to further investigate the direction of causality between these two variables using more advanced methods. The purpose of this paper is to reexamine the causality between exports and economic growth for the U.S. in a multivariate framework by employing Hsiao's version of the Granger causality method with the aid of cointegration and error-correction modeling. The remainder of the paper is organized as follows. Section II presents the methodology and model. Sections III and IV report the empirical results and conclusions.

II. METHODOLOGY AND MODEL

According to Granger (1969 and 1980), if some series x_t contains information in past terms that helps in the prediction of y_t and if this information is contained in no other series used in the prediction, then x_t is said to cause y_t. Thus, a variable x_t is said to Granger-cause y_t if the prediction of the current value of y_t is improved by using past values of x_t.

The Granger causality test method is chosen in this paper over alternative techniques such as the Sims test (Sims, 1972) and the Geweke test (Geweke, 1984) in light of the favorable Monte Carlo evidence reported by Guilkey and Salemi (1982), and Geweke, Meese and Dent (1983), particularly for small samples in applied work. Moreover, the Granger testing procedure can be readily generalized from a bivariate to a multivariate system.

(1) The Model

Following earlier studies (Michaely, 1977; Tyler, 1981; Ghartey, 1993), the role of exports in economic growth is analyzed using a standard production-function model which treats exports as a factor of production. Thus, the aggregate production function can be transformed into one error-correction term inclusive causality equation and rewritten by specifying a vector autoregressive model as follows:

$$\text{Log } Y = f(\log K, \log L, \log X), \tag{1}$$

where Y = GNP or income in constant dollars,

K = capital (CAP) in constant dollars,

L = labor (LAB),

X = exports (EXP) in constant dollars.

In testing for causality between exports and economic growth, a multivariate, rather than a bivariate model, is adopted in this study. While Ghartey (1993) applies a multivariate approach in his study on exports and economic growth for Japan, he employs a bivariate method for the United States. Bivariate casualty tests used in previous studies (e.g., Chow, 1987; Jung and Marshall, 1991; Bhamanni-Oskooee and Alse, 1993; Dutt and Ghosh, 1996 and others) have fallen out of favor in macroeconomics. For instance, Granger (1969), Sims (1980), Lutkepohl (1982) and Serletis (1988) have all demonstrated that Granger casualty is severely affected by the biases due to the omission of relevant variables. Further, standard production function theory suggests the potential importance of capital and labor. Therefore, these two additional variables are included in the exports/growth equation.

While employing a multivariate approach in his Japanese study, Ghartey (1993) did not perform any cointegration test and did not pay attention to the order in which the regressors were entered into the equation. Without a pretest of the cointegration property, there is no way to know whether the simple Granger test is appropriate. Consequently the causality inferences may be invalid. Further, failure to pay attention to the sequence in which each variable enters into the equation may produce erroneous causality results. The main contribution of this paper is methodological in that we apply the multivariate approach used by Ghartey in his study of Japan (Ghartey, 1993) with some modifications to reexamine the same issue for the United States. Such modifications include the performance of the Johansen cointegration test with the aid of an error-correction model which are not employed in Ghartey's study. In addition, unlike Ghartey's approach, the sequence in which each variable enters the estimating equation is determined by the specific gravity criterion (SGC) proposed by Caines, Keng, and Sethi (1981), as discussed below.

(2) Stationarity, Cointegration and Error-Correction Modeling

Hsiao's version of Granger causality test requires all variables to be stationary. The meaning of stationarity has been briefly discussed in the previous section. If economic data are stationary, they can be thought of as "nicely behaved." However, most economic data are trending. That is, the mean and variance change over time and thus clearly cannot be stationary. The difficulty that arises when employing standard regression techniques with clearly non-stationary series is that of the spurious correlation problem (Granger and Newbold, 1974)—this occurs when a correlation is found between two variables that are not causally related in theory. High R^2s may only indicate correlated trends and not true economic relationships. In such situations, as indicated previously, the usual ordinary least squares significance tests performed on the regression coefficients can be quite misleading. The Phillips-Perron (PP) test (Phillips and Perron, 1988) is the formal test for unit roots and stationarity used in this study. The PP test is robust for a variety of serial correlation and time-dependent hetroscedasticity problems.

As indicated, the concept of cointegration can be defined as a systematic co-movement among economic variables over the long run. Technically, according to Engle and Granger (1987), y_t, k_t, l_t, and x_t could each be I(1), yet, a particular combination of these four variables, $e_t = y_t - bk_t - cl_t - dx_t$, could still be I(0). Thus, if this property holds, y_t, k_t, l_t, and x_t are cointegrated.

If y_t, k_t, l_t and x_t are cointegrated, any standard Granger causal inference will be invalid and error-correction modeling should be incorporated into the system (Granger, 1988; Bhamani-Oskooee and Alse, 1993). Thus, a cointegration (CI) test is necessary before performing the Granger causality test. The Johansen cointegration test (Johansen, 1991; Johansen and Juselius, 1990) is performed. The Johansen cointegration test is chosen over the Engle-Granger two-step residual-based test because the former is a more robust cointegration test, particularly when a multivariate model is used. Johansen's procedure, which is based on full system estimation, can eliminate the simultaneous equation bias more efficiently than the Engle-Granger single equation method. Apart from this, Johansen's procedure tests for all possible cointegrating relations and does not suffer from problems associated with normalization—that is, from deciding which variable should be used as a dependent variable.

(3) Hsiao's Version of the Granger Multivariate Causality Method

While the Granger causality method discussed above is widely used in applied research, its application is unfortunately restricted to models where all variables entering the system have identical lag lengths. This assumption is not valid generally for macroeconomic time series data. Further, there is no theoretical reason why two or more variables have to have identical lag length. To overcome the shortcomings mentioned above, Hsiao (1981) suggests the following stepwise procedure for system identification. Equations (1) can be transformed into four autoregressive equations as follows:

$$(2) \qquad (1-L)y_t = \alpha_0 + \delta e_{t-1} + \sum_{i=1}^{M} \alpha_i (1-L) y_{t-i} + V_{1t},$$

$$(3) \qquad (1-L)y_t = \alpha_0 + \delta e_{t-1} + \sum_{i=1}^{M} \alpha_i (1-L) y_{t-i} + \sum_{j=1}^{N} \beta_j (1-L) x_{t-j} + V_{2t},$$

$$(4) \qquad (1-L)y_t = \alpha_0 + \delta e_{t-1} + \sum_{i=1}^{M} \alpha_i (1-L) y_{t-i} + \sum_{j=1}^{N} \beta_j (1-L) x_{t-j} + \sum_{k=1}^{P} \gamma_k (1-L) k_{t-k} + V_{3t},$$

$$(5) \qquad (1-L)y_t = \alpha_0 + \delta e_{t-1} + \sum_{i=1}^{M} \alpha_i (1-L) y_{t-i} + \sum_{j=1}^{N} \beta_j (1-L) x_{t-j} + \sum_{k=1}^{P} \eta_k (1-L) k_{t-k} + + \sum_{h=1}^{Q} \Phi_h (1-L) l_{t-h} + V_{4t},$$

where L is the lag operator, (1-L) = d is the difference operator such that $(1\text{-}L)y_t = y_t - y_{t-1}$ represents the first difference, v_{it} are the white noise disturbance terms and e_t are the stationary residuals from the multi-cointegration equation below (note that a sequence $\{v_t\}$ is a white-noise process if each value in the sequence has a mean of zero, a constant variance, and is serially uncorrelated):

$$(6) \qquad y_t = \alpha_0 + \alpha_1 x_t + \alpha_2 k_t + \alpha_3 l_t + e_t.$$

Using the univariate equation (2) for illustration, in step one we treat the dependent variable, y_t, as a one-dimensional autoregressive process and compute the sum of squared errors (SSE) using the same equation with the maximum order of lags set from 1 to M. The corresponding final prediction error (FPE), as defined by Akaike (1969), is calculated by using the following equation:

(7) $$\text{FPE}(m) = \frac{(T+m+1)}{(T-m-1)} \bullet \frac{SSE}{T},$$

where T = total number of observations,
m = the order of lags varying from 1 to M, and
SSE = the sum of squared errors.

Then, we choose the order which yields the smallest FPE, m*. In step two, focusing on the bivariate equation (3), we then treat y_t as a controlled variable, with the order of lags set at m*, and x_t as a manipulated variable. Using equations (3) and (8), we again compute the SSE and FPE of y_t by varying the order of lags of x_t from 1 to N and determine the order which yields the smallest FPE, n*. The corresponding two-dimensional FPE is

(8) $$\text{FPE}(m^*,n) = \frac{(T+m^*+n+1)}{(T-m^*-n-1)} \bullet \frac{SSE(m^*,n)}{T},$$

where n = the order of lags on x(t) varying from 1 to N.
m* = the optimum number of lags computed from (7).

If FPE(m*,n*) is less than FPE(m*), we then conclude that exports (x_t) Granger-cause economic growth (y_t). Subsequently, by using the same procedure, FPE(m*,n*) and FPE(m*, n*, p*) can be obtained and compared with each other. By repeating the same procedure for the export equations, causality from economic growth to exports can also be estimated.

Hsiao's method works well for a bivariate model. For a multivariate equation model, as indicated in previous section, the sequence in which each variable enters the estimating equation is determined by the specific gravity criterion (SGC) proposed by Caines, Keng, and Sethi (1981).

III. DATA AND EMPIRICAL RESULTS

This study covers the period 1940-90 for the United States. Annual data on GNP, exports, GNP deflators, capital and labor used in this study are compiled from the *Economic Report of the President* (1991 and 1994).

(1) Results of the Stationarity and Cointegration Tests

The results (Table 1) of the PP test indicate that GNP, exports, capital and labor individually are I(1). However, the PP test indicate that GNP, exports, capital and labor become stationary after first differencing.

The Johansen cointegration tests are conducted with 2, 4, 6, and 8 lags. As shown in Table 2, the cointegration tests reject the null hypothesis of no cointegrating vectors and support the null hypothesis of one or few cointegrating vectors. Cointegrating vectors are constraints that an economic system imposes on the movement of the variables in the long run. Therefore more cointegrating vectors mean a more stable system. Thus, the cointegration tests indicate that the four variables are cointegrated, and therefore that error-correction modelling must be adopted, as shown in equations (2) through (5).

(2) Results of the Granger Multivariate Causality Tests

Finally, after transforming the data, we perform the causality test. As stated earlier, if FPE(m*,n*) < FPE(m*), then exports Granger-cause GNP and vice versa. As shown in Table 3, the results reveal that for the GNP-export equation, the condition is met and therefore we conclude that exports Granger-cause GNP. When labor is added to the equation, the condition is again satisfied. We therefore infer that exports

TABLE 1. RESULTS OF THE UNIT ROOT AND COINTEGRATION TESTS

Variable	Phillips-Perron (PP) Tests	Critical Value
GNP	- 2.2109*	3.13
Exports	- 0.5289*	3.13
Capital	- 0.2135*	3.13
Labor	- 0.1834*	3.13

Note that all variables with the exception of labor are expressed in real terms.
*denotes non-stationary (for the critical value, see Fuller, 1976).

TABLE 2. RESULTS OF JOHANSEN COINTEGRATION TEST (TRACE TEST)

	LR Tests				
Rank (r)	Lag=2	Lag=4	Lag=6	Lag=8	Critical Value 99%
r=0	98.475*	95.277*	63.876*	2744.104*	60.054
r≤1	27.769	41.629*	29.606	1230.271*	40.198
r≤2	2.525	13.995	10.383	71.075*	24.988
r≤3	0.012	3.137	0.696	12.751*	12.741

*denotes significant at the 1% level (see Johansen and Juselius, 1990, p. 209, Table A3).

TABLE 3. RESULTS OF HSIAO'S VERSION OF GRANGER CAUSALITY TESTS AND THE OPTIMUM LAG OF THE VARIABLES

Controlled variable	First manipulated variable	Second manipulated variable	Third manipulated variable	FPE	Causality Inferences
Y(i=6)				0.7173099E-03	
Y(i=6)	X(j=3)			0.6658479E-03	X⇒Y
Y(i=6)	X(j=3)	L(m=3)		0.6462363E-03	L⇒Y
Y(i=6)	X(j=3)	L(m=3)	k(n=1)	0.6750432E-03	K≠>Y
X(i=8)				0.7084065E-02	
X(i=8)	Y(j=4)			0.6201597E-02	Y⇒X
X(i=8)	Y(j=4)	K(m=2)		0.5708027E-02	K⇒X
X(i=8)	Y(j=4)	K(m=2)	L(q=1)	0.5262660E-02	L⇒X

Note that Y denotes real GNP, X real exports, K real capital and L labor.
The figures in parentheses behind the variables are the optimal lag length.

and labor Granger-cause GNP. Subsequently, capital is added to the equation, and since FPE (m*,n*) > FPE (m*), we conclude that capital does not Granger-cause GNP. Nevertheless, following Hsiao (1982) if capital causes exports and exports cause GNP, then capital indirectly causes GNP. Conversely, for the export equation, the results in table 3 indicate that GNP Granger-causes exports. As such, our results show that GNP along with capital Granger-causes exports. Subsequently, labor is added to the equation and again we conclude that GNP, capital, and labor Granger-cause exports.

(3) Discussion of the Empirical Results

The study finds that a bidirectional causality between exports and economic growth exists along with capital and labor. Our findings on the feedback relationship between exports and economic growth are consistent with several multivariate cross-sectional studies for the U. S. (Kavoussi, 1984; Moschos, 1989 and others). Exports Granger-cause economic growth and exports are being Granger-caused by GNP. Our results, however, are not in complete agreement with the bivariate study by Ghartey (1993) in which economic growth is found to unidirectionally cause export growth in the United States without feedback. This indicates that an omission of relevant variables can cause biased causality results.

The finding of the bidirectional causality between exports and economic growth basically confirms the hypothesis that for a large developed country, export growth propels economic growth and in turn economic growth boosts export growth. Our finding that export growth causes economic growth is fully consistent with the conclusions of traditional trade theories as well as some recently developed theories (Krueger, 1980; Kunst and Marin, 1989; Lucas, 1990; Ghartey, 1993; Dutt and Ghosh, 1996). As indicated, the expected gains from exports according to these theories consist of improved allocative efficiency, scale economies, enhanced productivity, technological progress, and dynamic spill-over effects from the export sector to other sectors. In addition, according to neoclassical growth theory we would expect growth of these factors to cause economic growth.

The finding that labor growth causes economic growth also conforms to our expectations. Labor is heavily embodied with education and training in the United States. Our continuous investment and improvement in healthcare, education, training, and research have enhanced and upgraded the quality of our labor force considerably over the years. Although wage rates in the United States have risen rapidly compared with our trade partners, the productivity of labor has also increased considerably. Consequently, we still maintain a comparative advantage in certain areas.

The finding that capital affects GNP, filtered through exports, is interesting and revealing. In the literature, the accelerator theory (or supply-side view) holds that economic growth causes investment growth, while the multiplier theory (or the demand-side view) maintains that economic growth is a function of investment. No causal relationship between capital and GNP is found for Japan by Ghartey (1993).

The causal linkages of economic growth to exports have not been fully explored and addressed in the literature. The Linder thesis (Linder, 1961) states that rich nations tend to trade extensively with each other. The Linder hypothesis is further extended by Lancaster (1980) who emphasizes the phenomenon of two-way trade within the same industry. Kunst and Marin (1989) note that an improvement in productivity which is induced by the realization of economies of scale may give rise to exports. Additionally, if technical progress leads to well-developed markets, it may in turn improve export performance as well.

Further, the causal link from capital and labor growth to export growth, as reported in this study, to my knowledge, has been not been fully addressed in the empirical literature. Additional investment in education, technology, R & D, and the upgrading of aged equipment can improve productivity and make the country more competitive in international markets, thereby improving exports. An investment in education will also enhance workers' skills which in turn raise labor productivity and exports.

The results of this study also demonstrate that U.S.'s historical commitment to an outward-looking, open-economy policy has contributed to our strong economy today. Free trade can promote economic growth and economic welfare not only by improving efficiency through realization of a nation's comparative advantage, which is a movement along the production possibilities frontier from the equilibrium point in isolation to a new equilibrium point in an open economy, but also by stimulating growth, which is an outward movement of the production possibilities frontier. Therefore, the U.S. should continue its commitment to a free trade policy.

IV. CONCLUSIONS

This study finds that exports and labor influence economic growth and that past income, capital, and labor affect exports. This may indicate that export growth causes economic growth and this growth effect then enhances export performance through technical and spinoff effects. In other words, strong economic growth can make the economy more competitive which in turn promotes export growth and thereby increases the national income of the country. Thus, economic growth and export expansion could strengthen and reinforce each other in the growth process.

The results of this study generally support the hypothesis that for a large developed country, exports pull growth and growth pushes exports. The policy implications of this are that in general free trade is beneficial to a large developed country, such as the United States. A free trade policy promotes faster economic growth in the United States and a strong American economy lifts up the world economy as well.

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THE LONG-RUN RELATION BETWEEN U.S. STOCK MUTUAL FUNDS AND COMMERCIAL BANK TIME DEPOSITS: EVIDENCE FROM COINTEGRATION

Matiur Rahman* and Muhammad Mustafa**

I. INTRODUCTION

The explosive growth in stock mutual funds, which has already changed financial markets, is now transforming the economy itself. As a result, commercial banks may be the biggest losers. Their share of the U.S. credit market has declined to about 15 percent. Borrowers that once might have gone to banks are going to securities markets. This is changing the structure of the financial system. This might be due to the fact that households are increasingly becoming risk-takers, shifting their savings into stock mutual funds. From 1980 to mid-1993, mutual fund assets grew more than 12 times. Mutual funds provided 25 percent of the resources raised by domestic nonfinancial sectors as compared to 10 percent provided by depository institutions.

The time series data on financial variables (like macroeconomic variables) are very likely to be nonstationary, although studies in the early 1970s and prior decades assumed stationarity. Such a presumption was too simplistic and often misleading. To correct for this problem, the cointegration technique is applied in this paper. Most previous studies have concentrated primarily on mutual fund performance in isolation [i.e., Arditti, 1971; Carlson, 1970; Cohen and Pogue, 1968; Fabozzi, et al., 1980; Joy and Porter, 1974; Kon and Jen, 1978; McDonald, 1974; Meyer, 1977; Sharpe, 1966; Simonson, 1972; Treynor and Mazuy, 1966; West, 1968; and Williamson, 1972].

The primary objective of this paper is to examine the long-run relationship between the real volume of investment in stock mutual funds and the real volume of investment in time deposits with commercial banks. The remainder of the paper is organized as follows: Section II outlines the cointegration methodology. Section III reports the empirical results, analyzes them and offers some concluding remarks.

II. THE COINTEGRATION METHODOLOGY

The cointegration approach that has been applied in this paper is briefly outlined as follows:

(1) $$x_t = a_0 + a_1 y_t + z_t$$

where x_t = real volume of investment in stock mutual funds, y_t = real volume of investment in commercial bank time deposits, and z_t is the stochastic error term. The variables, x_t and y_t are cointegrated of order d (i.e., I(d)) if the time series data on them have to be differenced d times to restore stationarity. For d=0, x_t and y_t are stationary in levels and no differencing is necessary. For d=1, first differencing is needed to restore stationarity.

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First, following Engle and Granger [1987] the time series property of each variable is examined by unit root tests. For unit root tests, the following equations are considered:

(2) $$x_t = \mu + \beta T + \alpha x_{t-1} + \sum_{i=1}^{k} c_i \Delta x_{t-i}$$

(3) $$y_t = \theta + \pi T + \psi y_{t-1} + \sum_{i=1}^{k} d_i \Delta y_{t-i}$$

Each time series has non-zero mean and non-zero drift. That is why the estimation should include both a constant term (μ or θ) and a trend term (T) in each specification. The relevant null hypothesis is that $|\alpha| = 1$ or $|\psi| = 1$ against the corresponding alternative hypothesis that $|\alpha| < 1$ or $|\psi| < 1$. A failure to reject the null hypothesis would imply that each time series is nonstationary and stationarity can be restored by the first or higher order differencing of the level data. Two nonstationary time series have a potential for long-run convergence, if they are cointegrated.

To search further for cointegration, the following ADF regression that corresponds to cointegration regression (1) is considered:

(4) $$\Delta z_t = \alpha z_{t-1} + \sum_{i=1}^{m} b_i \Delta z_{t-i} + q_t$$

The ADF test is applied on â to accept or reject the null hypothesis of no-cointegration. The null hypothesis is rejected if the calculated pseudo t-value associated with $\hat{\alpha}$ is greater than its critical value, as provided in Engle and Yoo [1987] at various levels of significance.

If x_t and y_t are cointegrated, there must exist an error-correction representation which may take the following form:

(5) $$\Delta x_t = \beta_1 e_{t-1} + \sum_{i=1}^{k} \phi_i \Delta x_{t-i} + \sum_{j=1}^{k} \delta_j \Delta y_{t-j} + u_{1t}$$

If $\hat{\alpha}_1$ can be obtained from equation (1) so that z_t can be cointegrated, the remaining parameters in equation (5) can easily be estimated. Then the usual t-test can be applied to determine short-run and/or long-run linkages between x_t and y_t. The series x_t and y_t are cointegrated when $\hat{\beta}_1$ is significantly different from zero. The non-zero $\hat{\beta}_1$ captures the short-run influence of long-run dynamics. If $\hat{\beta}_1 \neq 0$, then y_t will lead x_t in the long run. If the $\hat{\delta}_j$ s are not all zero, movements in y_t will lead those in x_t in the short run.

The error-correction model (ECM) was first introduced by Sargan [1964] and subsequently popularized by Davidson et al. [1978] and Hendry et al. [1984], among others. It has enjoyed a revival in popularity due to the recent work of Granger [1986, 1988], and Engle and Granger [1987] on cointegration. Its importance in the cointegration literature derives from the fact that if two variables are cointegrated of order 1, they can be modeled as having been generated by an ECM.

Monthly data have been used in this study from 1966 through 1995 to divide this entire sample period into two equal sub-periods to represent pre- and post-deregulation eras. The pre-deregulation era is represented by 1966-1980 and the post-deregulation era is represented by 1981-1995. The data have been collected from *Standard and Poor's Basic Statistics*, the *Economic Report of the President*, and the *Federal Reserve Bulletin*.

III. EMPIRICAL RESULTS, ANALYSES AND CONCLUDING REMARKS

The ADF statistics (with and without trend) for unit root tests have been obtained by estimating equations (2) and (3). They are reported as follows:

TABLE 1. ADF TEST OF UNIT ROOT*

Variable	ADF (with trend)	ADF (without trend)	Optimum Number of Lags
	1966-1980		
VEM	-2.1955	-1.4634	4
VCD	-0.93063	-2.4512	4
	1981-1995		
VEM	-0.6746	-1.1377	4
VCD	-0.6880	-2.2394	4

*Critical value at 5% level of significance is -3.410 (with trend) and that without trend is -2.8600. Here, VEM = Real Volume of Investment in Equity Mutual Funds ($ million), and VCD = Real Volume of Investment in Commercial Bank Time Deposits ($ million). The optimum number of lags is determined by the final prediction error (FPE) criterion.

A comparison of the ADF values (with and without trend) against the respective critical values show that the null hypothesis of unit root (nonstationarity) cannot be rejected at the 5 percent level of significance. It implies that the real volume of investment in equity mutual funds and the real volume of investment in time deposits with commercial banks are non-stationary in the pre-deregulation and the post-deregulation eras.

Next, the ADF regression (4) is estimated to examine whether the two variables under consideration are cointegrated. The empirical results are provided as follows:

TABLE 2. COINTEGRATION TESTS BASED ON ADF PROCEDURES*

X_t	Y_t	ADF Statistics
	1966-1980	
VEM	VCD	-3.03 (5)
	1981-1995	
VEM	VCD	-3.953 (5)

* The critical values of ADF statistics, reported in Engle and Yoo [1987] are -4.07, -3.37, and -3.03 at 1, 5 and 10 percent levels of significance respectively. Lag-lengths are reported within parentheses. They are selected, based on final prediction error (FPE) criterion.

It is obvious that the null hypothesis of no-cointegration can be rejected only at 10 percent level of significance for both pre-deregulation and post-deregulation eras. In both cases, the relation is negative. Both the pre-deregulation and post-deregulation eras depict a long-run association between the real volume of investment in equity mutual funds and the real volume of investment in time deposits with commercial banks. As expected, the long-run negative relation is even stronger during the post-deregulation era than during the pre-deregulation era.

Finally, the estimates of the error-correction model (5) are provided as follows:

TABLE 3. BIVARIATE GRANGER-CAUSALITY TEST*

Dependent Variable	"Causal" Variable	Lag Orders	F-Statistic	$\bar{R}^2$
		1966-1980		
VEM	VCD	m=2, n=2	(2,9) = 1.3070	0.5231
		1981-1995		
VEM	VCD	m=1, n=2	(1,9) = 3.3577**	0.5345

*a. Lag orders are selected based on the FPE criterion, m=lag length of variable x dependent, n = lag length of "causal variable".

b. The F-statistics (with degrees of freedom in parentheses) tests the joint null hypothesis that all the coefficients of the "causal variable" are simultaneously equal to zero.

**Significant at 5% level.

The joint F-test reveals that there is a long-run causal connection between the real volume of investment in time deposits with commercial banks and the real volume of investment in stock mutual funds. More precisely, a decline in time deposits with commercial banks causes an increase in investment in stock mutual funds over time. In other words, the gain in the volume of stock mutual funds occurs at the expense of time deposits with commercial banks. There is no evidence of reverse causality. It implies that a fall in the volume of stock mutual funds does not cause a reverse flow of resources from the stock market to commercial banks in the form of time deposits. In such a case, resources will flow from the stock market to financial markets other than commercial banks.

Investment in stock mutual funds appears to be more attractive than investment in time deposits with commercial banks particularly in the post-deregulation era. Presumably, it is due to the fact that (i) stock mutual funds provide larger diversification because individual investors cannot diversify away security-specific risks to the same extent, (ii) stock mutual funds have lower transaction costs due to savings gained through brokerage and other security services discounts on large trades, and (iii) stock mutual funds enable investors to share liquidity risk through a commingling of investor assets.

These changes complicate the Federal Reserve's job of guarding against systematic risk. With capital flows faster and more liquid than ever, markets are more vulnerable to unexpected shocks. When commercial banks were dominant, the Fed could act as a lender of the last resort. Now, there is no such built-in safety valve. To survive, commercial banks are seeking laws and regulations allowing them into new businesses, such as insurance and stock underwriting. Some rule changes have already been enacted. As a result, banks will enjoy more product and market diversification. Commercial banks are now gearing up their activities in mutual fund sales. Currently, they enjoy a market share of 12 percent in aggregate mutual fund sales. To close, commercial banks need to undergo further transformation to be more viable in the increasingly fiercely competitive and dynamic capital markets.

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The Effects of Television Advertising on Concentration: An Update

Arthur S. Leahy*

This note presents an update of the empirical evidence on the relationship between television advertising and concentration. Since recent surveys of the evidence on the relationship between advertising and concentration have yielded vastly different conclusions, this issue is far from being resolved (Leahy, 1997; Tokle, 1993). Leahy's survey suggests that the empirical results vary according to the industries included in the sample, the methodology employed in the estimation of this relationship, and the period under consideration. Alternatively, the survey by Tokle concludes that higher levels of advertising have led to higher levels of concentration in the past. Furthermore, both Tokle and Leahy suggest that television advertising, in particular, has been isolated as having a distinct impact on concentration.

Three previous studies in this journal have obtained mixed results on the latter issue, although none of these studies has found a negative and significant effect of television advertising on concentration. In one of those studies, Leahy (1988) finds a positive and significant effect of television advertising on concentration for 1947-1967 and an insignificant effect for 1947-1972. In a later study, Leahy (1989) obtains a positive and significant effect of television advertising on concentration from 1947-1977. Tokle (1995) obtains a similar result for 1967-1982 and 1967-1987 for a larger sample of industries using advertising data for both 1967 and 1982. A possible problem with the latter study has been noted by Leahy (1988) in his criticism of Lynk (1981). Lynk's industry samples were from the period 1954-1963 and 1954-1967. Leahy argued that the choice of the time period examined by Lynk was likely to have resulted in the omission of a major portion of the structural impact of the introduction of television. The results of the Tokle study could be affected by this problem as well.

Regular network television service began in 1946. In 1947, the first regularly scheduled drama series and the first World Series were carried by a network. Nationwide television began in 1951 (Brooks and Marsh, 1981). By 1954, more than 55 percent of all households in the United States owned a television set (Lynk, 1981). There has not been much change in the structure of network television since the mid-1950s (Brooks and Marsh, 1981).

Using the methodology employed in previous studies by Leahy, the relationship between television advertising and concentration is examined over a longer period.[1] The following equation has been estimated (t statistics in parentheses):

$$\Delta CR = .32 - \underset{(3.19)}{.29}\,CR + \underset{(2.65)}{.14}\,TV - \underset{(1.54)}{.004}\,GR - \underset{(2.1)}{.04}\,S;$$

$$R^2 = .25,\ N = 62$$

where Δ CR and GR are the change in the four-firm concentration ratio and the industry growth rate, respectively, over the 1947- 1982 period; TV is the ratio of television advertising to total advertising in 1967; CR is the concentration ratio in 1947;[2] and S is industry size. The TV and CR variables are the same as those used in Leahy (1988), while Δ CR and GR are measured over the longer time period.

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Following Mueller and Rogers (1980), Leahy (1989), and Tokle (1995), industry size (S) is included in the equation. It is measured as the natural logarithm of industry value added in 1947.

As can be seen from the results above, the CR, GR, and S variables have the expected negative signs. In addition, the coefficients of the CR, TV, and S variables are statistically significant at the 5 percent level or better. Consistent with the findings of Mueller and Rogers, Leahy (1989), and Tokle (1995), both the advertising and industry size variables are significant determinants of changes in concentration over the period.

The above equation was also estimated over the 1947-1987 period. All variables are the same as those above except that Δ CR and GR are measured over the longer time period. The following results were obtained (t statistics in parentheses):

$$D\,CR = .34 - \underset{(3.12)}{.34\,CR} + \underset{(1.77)}{.11\,TV} - \underset{(.49)}{.0006\,GR} - \underset{(1.57)}{.03\,S};$$

$$R2 = .19,\ N = 57$$

The results are similar to those obtained previously, except that the coefficient of the TV variable has decreased in significance from the 1 percent to the 10 percent level and the coefficient of the S variable is no longer significant. The positive and significant coefficient of the TV variable in the above results suggests that the large-firm effect of advertising on concentration continues to outweigh the entry effect.[3] The results obtained here, therefore, provide more recent evidence regarding the relative magnitudes of the two effects of television advertising on concentration. Alternatively, as suggested by Tokle (1995), if television advertising increases concentration because it is a more effective means of product differentiation and creates barriers to entry, the results of this study support this conclusion.

ENDNOTES

1. Due to missing Census data, the following industries were deleted from those listed in Appendix A of the Leahy (1988) study: 2042, 2071, 2072, 2073, 2111, 2511, 2821, 2892, 3021, 3141, 3481, 3492, 3722, 3871, 3944, and 3953 for the 1947-1982 sample. For the 1947-1987 sample, the following additional industries were deleted: 2292, 2327, 3263, 3291, and 3964.
2. See Leahy (1988) for a discussion of a bias problem associated with the use of this variable.
3. See Leahy (1988) for a theoretical discussion of the large-firm effect and the entry effect.

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REFEREES

1. Vincent Benziger
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3. Patricia Chadwell
4. Dal Didia
5. Rick Fenner
6. Althea Hojnacki
7. Barbara Howard
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NEW YORK STATE ECONOMICS ASSOCIATION (NYSEA)

49th ANNUAL CONVENTION

FINAL PROGRAM

Triphammer Lodge – Ithaca, New York

September 27-28, 1996

Friday, September 27

8:00 – 10:00 PM	**NYSEA Convention Opening Reception** (Cortland Room)
	Wine and Cheese Reception
	Speaker: Bill Kolberg, Ithaca College *"The Use of Computers in the Classroom"—Demonstration and Open Discussion"*

Saturday, September 28

8:00 – 10:00 AM	**Convention Registration** (Cortland Room) Pick up final program, receipt/register, location directions, name tags
8:00 – 2:00 PM	**Textbook Display** (Cortland Room)
	Morning refreshments—Cortland Room
12:00 – 1:30 PM	**Luncheon** (Cortland Room)
	Speaker: Laurence Kotlikoff, Boston University *"American Savings and Fiscal Crises"*
	Afternoon Refreshments—Cortland Room
8:30 AM	**Sessions Begin**

SESSION I

8:30 – 10:00 AM **Public Finance** (Red Room)
Chair: Charles Callahan, III, SUNY Brockport

Dixie Blackley and Edward Shepard, LeMoyne College
"The Diffusion of Innovation in Home Building"
Discussant:
William O'Dea, SUNY Oneonta

Robert Jones, Skidmore College
"Sectoral Employment Cycles within the State and City of New York"
Discussant:
TBA

Kenneth Wickman
"Shift-Share Analysis of the Changing Economic Structure of Regions and Counties in New York"
Discussant:
John Page, Dominican College

SESSION II

8:30 –10:00 AM **Industrial Organization and Finance** (Green Room)
Chair: Dale Tussing, Syracuse University

Joseph Cheng and Abraham Mulugetta, Ithaca College
"Arbitrating Between the S&P 500 and Interest Rates"
Discussant:
Barbara Howard, SUNY Geneseo

Thomas Kopp, Siena College
"Acquisitions/Divestitures: A Case Study"
Discussant:
Elia Kacapyr, Ithaca College

Mehmet Karaaslan, Alfred University
"A Model of Technology Induced Merger"
Discussant:
Sherry Wetchler, Ithaca College

SESSION III

10:15 – 11:45 AM **The Quebec Economy: Contemporary Political, Environmental and Health Care Issues** (Corning Room)

Chair and Organizer:

Myrna Delson-Karan, Quebec Government House

Marie Agen, Syracuse University
"Projections on the Economic Impact of Quebec Sovereignty"

Ellen Fitzpatrick and Joy Wilson, SUNY Plattsburgh
"The Quebec/NY Environmental Stewardship of Lake Champlain"

Loretta Pagnotta, SUNY Plattsburgh
"The Economics of Rural Health Care in Quebec"

SESSION IV

10:15 – 11:45 AM **Macro** (Red Room)

Chair: James Booker, Alfred University

Elia Kacapyr, Ithaca College
"American Well-Being"
Discussant:
Thomas Kopp, Siena College

David Ring, Ithaca College
"Has the Growth of Potential Output Accelerated in the 1990s?"
Discussant:
Joseph Cheng, Siena College

SESSION V

10:15 – 11:45 AM **Economic Methodology** (Green Room)

Chair: Kent Klitgaard, Wells College

William Ganley, Buffalo State College
"A Veblenian Tale of Economic Methodology"
Discussant:
Kent Klitgaard, Wells College

SESSION VI

1:45 – 3:15 PM **International** (Red Room)

Chair: Alfred Lubell, SUNY Oneonta

John Humphries
"The Overseas Movement of American Currency"
Discussant:
Lynn Smith, Clarion University of Pennsylvania

SESSION VII

1:45 – 3:15 PM **Economic Education** (Green Room)

Chair: Barbara Howard, SUNY Geneseo

Charles O'Donnell and Joseph Ford, Iona College
"Redesigning the Economics Course in the MBA Core Curriculum: A Case Study"
Discussant:
Frank Musgrave, Ithaca College

John Page, Dominican College of Blauvelt
"College Economic Texts: Motivating or Intimidating"
Discussant:
David Pate, St. John Fisher College

SESSION VIII

1:45 – 3:15 PM **Environmental Economics** (Corning Room)

Chair: Dixie Blackley, LeMoyne College

Kent Klitgaard, Wells College
"Ecological Economies and Monopoly Capitalism: Questions Regarding Long-Term Sustainability"
Discussant:
Mehmet Karaaslan, Alfred University

James Booker, Alfred University
"Recreational Benefits at Allegheny State Park"
Discussant:
David Ring, SUNY Oneonta

William O'Dea, SUNY Oneonta
"Determination of the Optimal Speed Limit and Scale of a Highway"
Discussant:
Charles Callahan, III, SUNY Brockport

SESSION IX

3:30 – 5:00 PM **Results of Leftist Movements** (Red Room)

Chair and Organizer:

Bogdan Mieczkowski, Ithaca College

Po-Chih Lee
"Economic Effects of Communism in China"

Savo Jevremovic, Alfred University
"Inflation Problems in the Former Yugoslavia"

Bogdan Mieczkowski, Ithaca College
"The Economic Program and the Results of the New Left Movement of the 1960s and 1970s in the U.S."

Discussants:

Wade Thomas, SUNY Oneonta
Alfred Lubell, SUNY Oneonta
Hiroshi Ito, SUNY Plattsburgh

5:00 – 6:00 PM **NYSEA Business Meeting** (Corning Room)

NYSEA