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After-Hours Stock Prices and Post-Crash Hangovers

DAVID NEUMARK, P. A. TINSLEY, and SUZANNE TOSINI*

ABSTRACT

After-hours pricing in foreign equity markets of multiple-listed U.S. securities appeared to be efficient in predicting New York prices in the weeks immediately following the October 1987 crash but relatively uninformative in succeeding months. By contrast, daily changes in New York prices appear to be efficiently incorporated in after-hours trading on both the Tokyo and London exchanges throughout the sample period. This paper suggests that the asymmetry and temporal variations in cross-market correlations are consistent with rational investor behavior in equity markets with nonzero transaction costs and time-varying share price volatility.

In the wake of the October 1987 crash, a number of studies have provided evidence of significant international linkages among national equity markets where high-frequency movements in the share price indices of national exchanges appear to induce sympathetic price movements in subsequent trading on other national exchanges.[1] However, the cross-market correlations are generally much larger in periods of extreme volatility and appear to subside to modest or even negligible associations during periods of more normal trading activity. Since the constituent stocks of national exchange indices are not identical, it may be that the episodic increases in correlated movements of the indices are due to changes in the actual (or perceived) relative importance of common factors in periods of unusual volatility.

The thrust of this paper is to remove the issue of the disparate composition of the price indices of national exchanges by examining the prices of a set of

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[1] Recent studies include Bennett and Kelleher (1988), Bertero and Mayer (1989), Hamao, Masulis, and Ng (1990), and King and Wadhwani (1990). Evidence in these studies on the predictive content of international price movements may be contrasted with the Brady (1988) report's dismissal of the transmission role of international equity markets as a significant contributor to the events of October 1987. Although cross-correlations of movements in the aggregate price indices of national exchanges were examined in a number of pre-crash studies, share price comovements across exchanges were generally insignificant or unstable, such as in Agmon (1972), Bertoneche (1979), and Panton, Lessig, and Joy (1976). More robust associations among national exchange index movements were reported by Hilliard (1979) and Schollhammer and Sand (1985).

multiple-listed stocks as they are valued in successive trading on different national equity markets. Specifically, we examine the U.S. and foreign pricing of stocks in the Dow Jones index that are listed also on the London and/or Tokyo exchange. An interesting finding is that the same dynamic pattern of association is found in tests of the efficiency of cross-market pricing of homogeneous shares. That is, the cross-market responses are strong and not inconsistent with efficient market pricing in the immediate aftermath of the 1987 crash, but the U.S. responses to foreign pricing then dwindle to negligible levels, suggesting inefficient use of cross-market information in more normal trading periods.

Since this dynamic pattern of cross-market correlations cannot be ascribed to shifts in the relative importance of common factors, it is suggested that investor transaction costs provide a plausible explanation of variations in apparent cross-market efficiency. Because only larger price changes pierce the transaction cost barriers between markets, associations between markets will appear to be stronger in periods of unusually large price volatility. The root origins of episodes of extreme volatility are not explored here, but the empirical results presented below suggest that the joint interactions of investor transaction costs and share price volatility appear to have sizeable impacts on the international transmission of equity price movements.[2]

I. The Data

The basic data consist of daily prices of multiple listings of selected U.S. stocks on the New York, Tokyo, and London exchanges for the period October 21, 1987 through June 15, 1988. New York stock prices are for the thirty U.S. stocks that comprise the Dow Jones Industrial Average. Prices of most of the Dow Jones stocks traded on the Tokyo and London exchanges are used.[3] Tokyo data are closing exchange prices converted into U.S. dollars, while the London data, also expressed in U.S. dollars, are averages of bid-ask spreads rather than actual transaction prices.

The daily timing of trading on the three exchanges and the collection of sample price data are illustrated in the 24-hour "clock" of Figure 1. For the purpose of notation, the international trading day is defined to "begin" at 9:30 A.M. (EST) with the opening of the New York exchange. The collected data include the opening value of the Dow Jones index but not of individual stock prices. Moving sequentially through the international trading day,

[2] A recent empirical review of various determinants of equity price volatility is presented in Schwert (1989).

[3] The Tokyo listings are: American Express, DuPont, Eastman Kodak, Exxon, General Motors, IBM, McDonalds, Minnesota Mining, Phillip Morris, and Proctor and Gamble. The London listings are: American Express, DuPont, Eastman Kodak, Exxon, General Motors, IBM, AT&T, Boeing, Chevron, Coca Cola, General Electric, International Paper, Sears, and USX. Volume data for London are unavailable; volume data for Tokyo indicate that, in the period studied, the volume of trading in U.S. stocks listed on the Tokyo exchange was somewhat less than 10 percent of the volume in New York.

TRADING AND SAMPLING TIMES

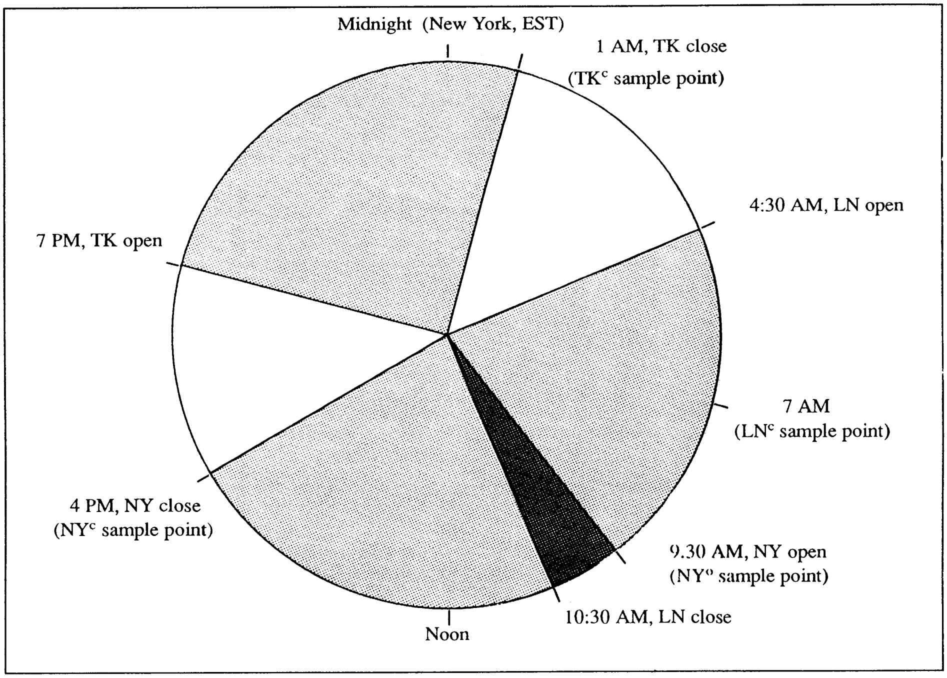

Figure 1. Trading and Sampling Times. Daily timing of trading on New York, Tokyo, and London exchanges and collection of sample price data at Tokyo close (TK^c), London 7 AM, EST (LN^c), New York open (NY^o), and New York close (NY^c).

data on individual share prices were collected for New York listings at the 4:00 P.M. closing, for Tokyo listings at the 1:00 A.M. (EST) closing, and for London listings at 7:00 A.M. (EST) (this will be referred to as the London "close").[4]

There are two sources of asynchronicity in the data. First, in the sample period studied, trading occurred on the Tokyo exchange on each Saturday

[4] There are some differences in collection procedures between initial and subsequent observations. Initially, Tokyo closing prices were obtained from the Bank of Japan and converted to dollars using the closing exchange rate; after February 22, closing yen prices and exchange rates were collected from Reuters about 3 hours after the Tokyo close. Prior to January 18, London prices, quoted in dollars by the Bank of England, were averages of the highest bid and lowest ask prices of all dealers as of 6:30 A.M.; subsequent London prices were averages of the bid and ask prices from First Boston as of 7:00 A.M.. Regression tests, allowing coefficients to be altered by changes in data collection procedures, rejected the possibility that changes in the timing and scope of the sample data affected the regression results below. In the cases of Tables I and II, for example, the apparent reductions in cross-market efficiency appear also in sample periods ending in mid-February (Tokyo) or mid-January (London).

except the second and third Saturday of the month. To maintain a consistent time structure of data across exchanges, Tokyo Saturday prices are not used. Information revealed in Saturday trading will be contained in the Friday-to-Monday Tokyo price changes transmitted to the Monday opening of the New York exchange (see Barclay, Litzenberger, and Warner (1988)). Second, there are a number of trading days with missing data due to national holidays and technical problems in data retrieval. This problem is resolved by forming all of the 1-day price changes used in the regression analyses and then dropping observations for days on which the transformed data are missing for at least one of the markets.

II. Regression Tests of the Efficiency of After-Hours Pricing

Under the efficient market hypothesis, high-frequency movements of the (log) price of a given U.S. equity share p_t may be characterized as a martingale process

$$p_t - p_{t'} = a(t - t'), \tag{1}$$

where the stochastic disturbance $a(t - t')$ denotes the impact of relevant "news" received during the time interval $t - t'$.

If all relevant markets are efficient, an after-hours price of the same U.S. share on a foreign exchange f_{t_*} should reflect the evolving value of the stock as it incorporates news received during the remaining hours of the international trading day

$$f_{t_*} = p_{t_*}, \tag{2}$$

where, of course, both prices are denominated in a common currency.

Efficiency in international equity markets implies that foreign after-hours price movements provide an unbiased prediction of overnight changes in U.S. home market prices. Thus, a conventional regression test of sequential pricing efficiency

$$p_t - p_{t'} = \alpha + \beta\left[f_{t_*} - p_{t'}\right] + a(t - t*), \qquad t' < t* \leq t, \tag{3}$$

may be expected to yield a unit slope coefficient $\beta = 1$.

Empirical estimates of equation (3) are presented below for prices of individual stocks and for approximate Dow Jones price indices for each exchange. The foreign Dow Jones indices measure the price (in U.S. dollars) of the Dow Jones index at the close of the relevant foreign market. The Tokyo Dow Jones index TK_t^c and the London Dow Jones LN_t^c are constructed from available listings of Dow Jones constituent stocks on the relevant foreign exchange. Using $NY_{i,t}^c$ to denote the closing price of stock i on day t in New York, NY_t^c to represent the closing of the Dow Jones index on day t, and $TK_{i,t}^c$ and $LN_{i,t}^c$ to indicate the subsequent price at the close of trading in the relevant foreign market, the foreign Dow Jones index constructions are

Table I

Regression Estimates of Dow Jones Overnight Price Response $\hat{\beta}$ to After-Hours Price Movements Using Daily Data from October 21, 1987 through June 15, 1988

The table provides full-sample and partitioned-sample estimates of the regression response $\hat{\beta}$ of overnight (close-to-open) percentage changes in the New York Dow Jones index $(NY_t^o - NY_{t-1}^c)/NY_{t-1}^c$ to after-hours (New York close-to-foreign close) percentage changes of Dow Jones indices on the Tokyo, $(TK_{t-1}^c - NY_{t-1}^c)/NY_{t-1}^c$, and London, $(LN_{t-1}^c - NY_{t-1}^c)/NY_{t-1}^c$, exchanges. The foreign Dow Jones index regressors are constructed from available Dow Jones share listings on the Tokyo and London exchanges. There are 110 daily observations for the Tokyo regressions, equations (1) to (3), and 135 daily observations for the London regressions, equations (4) to (6). Standard deviations of the regression coefficients are given in parentheses. Asterisks indicate violations of the efficient market hypothesis that the regression coefficients $\hat{\beta}$ are unity, using five percent critical values.

		Trading Days Following October 20, 1987							
	Intercept	Full Sample	Days 1-30	Days 30+	Days 1-10	Days 11-20	Days 20+	R^2	*DW*
Panel A: New York Dow Jones close-to-open response to Tokyo after-hours pricing									
(1)	-.001	.539	-	-	-	-	-	.095	1.36
	(.001)	(.161)*							
(2)	-.001	-	.934	.041	-	-	-	.158	1.10
	(.001)		(.209)	(.234)*					
(3)	-.001	-	-	-	.912	1.872	.094	.155	1.13
	(.001)				(.217)	(1.437)	(.229)*		
Panel B: New York Dow Jones close-to-open response to London after-hours pricing									
(4)	.002	.962	-	-	-	-	-	.364	1.78
	(.001)	(.110)							
(5)	.002	-	1.132	.478	-	-	-	.398	1.73
	(.001)		(.125)	(.208)*					
(6)	.001	-	-	-	1.633	.939	.408	.464	1.81
	(.001)				(.182)*	(.180)	(.166)*		

defined by

$$TK_t^c = \left[\frac{\sum_{i \in TK} TK_{i,t}^c}{\sum_{i \in TK} NY_{i,t}^c} \right] NY_t^c,$$

$$LN_t^c = \left[\frac{\sum_{i \in LN} LN_{i,t}^c}{\sum_{i \in LN} NY_{i,t}^c} \right] NY_t^c. \tag{4}$$

A. Initial Regressions

Table I reports results for regressions of overnight close-to-open movements in the U.S. Dow Jones index on after-hours movements of the Tokyo (Panel

A) and London (Panel B) Dow Jones index constructions.[5] Rejections of the hypothesis of efficiency $\beta = 1$ at the five percent level of significance are indicated by an asterisk. In general, a unit slope is maintained only for a brief interval following the October 1987 crash.

For example, in the second row of Table I, separate responses to after-hours Tokyo price movements are estimated for the first 30 trading days after the October collapse and the remaining sample observations. The results in the second row indicate that the mean New York response $\hat{\beta}$ is not significantly different from unity (0.934) in the 30 days immediately after the October collapse but that the estimated mean response in the remainder of the sample is not significantly different from zero (0.041). Thus, the efficiency of the New York and Tokyo markets appears to be maintained for 30 days following the October 1987 crash but rejected for trading days further removed from the collapse.

These interval snapshots of U.S. investor responses to after-hours foreign price movements are repeated in Panel B of Table I for the Dow Jones indicator constructed from London stock prices. As in the case of the Tokyo data, the fifth and sixth rows of Table I generally indicate that the estimated mean response coefficients of New York prices to overnight changes in the London prices are not significantly smaller than unity in the immediate aftermath of the October collapse but do fall significantly below unity in the observations beginning 20 and 30 trading days after the crash.

Table II presents estimates of similar specifications for individual prices of those Dow Jones stocks that trade in both Tokyo and London. In part because the dependent variable is now the close-to-close New York price movement, the results are not as statistically sharp as those reported for the Dow Jones indices. Nonetheless, the same qualitative results as in Table I are reflected in these estimates. For both Tokyo and London, the response coefficient for the first 10-day period is usually not significantly different from one (with the exceptions of the New York response to Tokyo prices for Kodak and General Motors). Generally, the estimated response in the second 10-day period is not significantly different from unity for the London prices but is significantly less than unity for the Tokyo prices, indicating a faster return to long-run transmission behavior for the Tokyo prices.

B. Competing Explanations

Before turning to interpretation of the apparent time variation in New York investor response to after-hours pricing, it is worth considering whether

[5] Prior to estimation of the regressions presented in this section, each series was checked for stationarity. At least at high-frequencies, stock prices (and stock price indices) are generally characterized as martingales or, in the case of i.i.d. disturbances, random walks. Dickey-Fuller tests indicated that unit roots were present in the daily New York, London, and Tokyo Dow Jones indices and most individual stock price series, and that first differences provided stationary series. Other variables that appear in the regressions are cross-market differentials between foreign and New York prices (or indices). These cross-market differentials also appeared to be stationary over the available sample, indicating the absence of persistent discrepancies in cross-market valuations.

Table II

Regression Estimates of Individual New York Share Price Responses $\hat{\beta}$ to After-Hours Price Movements Using Daily Data from October 21, 1987 through June 15, 1988

The table provides sample partitions of the regression response $\hat{\beta}$ of close-to-close percentage changes in an individual New York share price, $(NY^c_{i,t} - NY^c_{i,t-1})/NY^c_{i,t-1}$, to after-hours (New York close-to-foreign close) percentage changes in the price of corresponding foreign listings on the Tokyo, $(TK^c_{i,t-1} - NY^c_{i,t-1})/NY^c_{i,t-1}$, and London $(LN^c_{i,t-1} - NY^c_{i,t-1})/NY^c_{i,t-1}$, exchanges. Standard deviations of the regression coefficients are given in parentheses. Asterisks indicate violations of the efficient market hypothesis that the regression responses $\hat{\beta}$ are unity, using five percent critical values.

		Trading Days Following October 20, 1987					
	Intercept	Days 1-10	Days 11-20	Days 20+	N	R^2	DW
Panel A: New York close-to-close price response to Tokyo after-hours pricing							
American	−.001	1.129	−.314	−.139	139	.068	2.05
Express	(.003)	(.392)	(.389)*	(.175)*			
DuPont	.000	1.248	−.021	.190	143	.091	1.79
	(.002)	(.355)	(.322)*	(.155)*			
Kodak	−.003	.289	.229	.437	143	.049	1.87
	(.002)	(.192)*	(.452)	(.196)*			
Exxon	−.001	.791	−.334	.328	118	.084	1.92
	(.002)	(.383)	(.332)*	(.143)*			
General	.000	.158	.191	.359	139	.038	1.98
Motors	(.002)	(.182)*	(.249)*	(.169)*			
IBM	−.002	.909	−.314	.408	143	.188	2.09
	(.002)	(.168)	(.575)*	(.239)*			
Panel B: New York close-to-close price response to London after-hours pricing							
American	.000	1.406	.613	.177	139	.277	2.01
Express	(.002)	(.203)	(.371)	(.167)*			
DuPont	.001	1.056	1.389	.931	138	.184	1.86
	(.002)	(.239)	(.632)	(.373)			
Kodak	.003	.867	1.050	.878	139	.026	2.78
	(.006)	(.858)	(1.175)	(.634)			
Exxon	.002	1.469	1.179	.634	139	.185	1.91
	(.002)	(.316)	(.521)	(.306)			
General	.002	1.079	.169	.652	139	.100	2.02
Motors	(.002)	(.310)	(.371)*	(.385)			
IBM	−.000	1.126	−.013	−.070	139	.178	2.05
	(.001)	(.208)	(.419)*	(.326)*			

these results are a result of anomalies in the data used or of unusual institutional features of the particular foreign exchanges.

First, for the London market, averages of bid-ask spreads were used rather than actual transaction prices. Since trading volume in U.S. stocks on foreign markets was significantly higher in the period immediately following the October collapse, it is possible that these bid-ask averages were more accurate estimates of transaction prices in this period and became noisier as volume fell in periods further from the collapse. By a simple measurement error argument this would lead to declining estimates of β. However, the price data for Tokyo are actual transaction prices, so a measurement error explanation cannot explain the parallel (and, indeed, statistically stronger) finding of declining $\hat{\beta}$ values for the Tokyo market. Further, it is likely that trading imbalances that occurred in the period immediately following the collapse increased the noise content of bid-ask averages despite the higher volume of trading.[6]

Second, a much-discussed institutional feature of the Tokyo market is the imposition of daily limits on share price movements. The actual limits for any day depend on the previous day's closing price. On average, the price movement limit amounts to approximately 11–15 percent of this closing price.[7] However, the existence of these limits cannot explain the results in Tables I and II. While the TOPIX (a comprehensive index of all stocks in the "1st section") fell 15 percent on October 20, 1987, the data used here begin on October 21, 1987. Daily price limits were not binding during the sample period: the largest change (in absolute value) of the Tokyo Dow Jones index was 4.7 percent, while for individual stocks the largest change was 8.6 percent.

Third, in the case of Table I, use of the foreign Dow Jones indices, defined in equation (4), may introduce a systematic bias against the hypothesis of efficient market responses $\beta = 1$ because the foreign index regressors are based on price movements of only those subsets of Dow Jones stocks that are traded on the foreign exchanges. Thus, a percentage change in a constructed foreign subindex, say $\Delta \hat{f}_t$, may be a biased indicator of the concurrent expected percentage change in the full Dow Jones index Δf_t, viz.

$$\Delta f_t = \gamma \Delta \hat{f}_t + e_t, \tag{5}$$

where e_t denotes the prediction error and $(1 - \gamma)$ is a measure of the bias associated with use of the subindex $\hat{f}$. By this interpretation, the reductions of the estimated regression responses in Table I may be attributed to sample increases in the estimated bias of the foreign subindices $1 - \gamma$.[8] Although

[6] Asymmetric spreads or lags in posting of bid-ask spreads behind actual transaction prices might explain the estimated hypersensitive response to London price changes in the first 10 days after the crash.

[7] The *Tokyo Stock Exchange Fact Book* (1988) reports that the average price of stocks in 1987 in the "1st section" (representing 96.7 percent of capitalization on the Tokyo exchange) ranged from a low of 926 yen to a high of 1,300 yen. According to the daily price limit schedule, this implies average price movement limits from 11–15 percent. For foreign stocks, the relevant base price is the closing price on the home or other major overseas markets. Therefore, if the prices of foreign listings change in intervening markets, the close-to-close change on the Tokyo market may exceed the daily price limit.

[8] Suppose the price change of each share in the Dow Jones index consists of changes due to weighted common factors and changes due to share-specific factors. The residual e_t of equation (5) captures the impact of share-specific factors excluded from the subindex $\hat{f}$. The common factors of share prices included in the subindex are positively correlated with the common factors of share prices excluded from the subindex but included in the full index f. Consequently, the estimated slope coefficient γ of the subindex regressor in (5) will generally be an increasing function of the relative importance of the common factors and a decreasing function of the relative importance of those share-specific factors that are included in the subindex.

Table III

Regression Tests of the Indicator Bias of the Tokyo and London Dow Jones Index Constructions Using Daily Data from October 21, 1987 through June 15, 1988

The table presents estimates of the indicator bias that may be associated with the foreign Dow Jones indices constructed by the authors. Because only subsets of the stocks in the Dow Jones index are listed on the Tokyo and London exchanges, changes in the subset indices based on foreign listings may provide biased predictions of the expected change in the full Dow Jones index. The table lists the slope coefficients generated by regressing changes in the full Dow Jones index on concurrent changes in selected subset indices. In each regression, the dependent variable is the New York percentage change (close-to-close) of the Dow Jones index. The regressors are sample partitions of the New York percentage change (close-to-close) of the price average of those Dow Jones shares that are listed either on the Tokyo exchange, equations (1) and (2), or on the London exchange, equations (3) and (4). The indicator bias of a subindex is measured by one minus the estimated regression coefficient of the subindex; the standard deviation of each estimated regression coefficient is enclosed by parentheses.

		Trading Days Following October 20, 1987							
Regressors	Intercept	Days 1–30	Days 30+	Days 1–10	Days 11–20	Days 20+	N	R^2	DW
Tokyo Index									
(1)	.000	1.058	.868	–	–	–	104	.93	1.17
	(.001)	(.045)	(.032)						
(2)	.000	–	–	1.065	1.033	.882	104	.92	1.13
	(.001)			(.057)	(.135)	(.031)			
London Index									
(3)	.000	1.001	.805	–	–	–	129	.89	2.45
	(.001)	(.044)	(.035)						
(4)	.000	–	–	.994	1.038	.821	129	.89	2.40
	(.001)			(.056)	(.103)	(.034)			

foreign after-hours changes of the full Dow Jones index Δf are unobserved, representative estimates of the sample indicator bias that is due to use of selected subsets of shares can be based on the price movements of New York listings. Table III presents regressions of the New York close-to-close movements of the full Dow Jones index on the New York close-to-close price movements of the subsets of shares incorporated in the Tokyo and London subindices. In the immediate vicinity of the October 1987 crash, the estimated slope coefficients $\hat{\gamma}$ of the subset indices are not significantly different from unity. The estimated regression coefficients do fall in the remainder of the sample, suggesting an indicator bias $1 - \gamma$ of about 10 percent for the Tokyo index and about 20 percent for the London index. However, these systematic biases are too small to account for the reductions in estimated responses of 50–90 percent presented in Table I.

Thus, the principal conclusion that emerges from simple regression analysis of the relationship between foreign and domestic price movements is that the evidence is generally inconsistent with the efficient markets hypothesis. After-hours foreign price movements of U.S. stocks do not appear to be fully transmitted to opening New York prices. However, this finding is reversed for both Tokyo and London prices in relatively brief intervals immediately following the October 1987 collapse. The next section explores a transaction cost interpretation of the empirical finding that after-hours foreign pricing of U.S. stocks appears to be efficient only in turbulent periods.

III. A Transaction Cost Interpretation and Confirming Evidence

This section explores the possibility that the estimated time variation in apparent cross-market efficiency is not an anomaly but a consequence of rational investor behavior.

In the absence of transaction costs, it is presumed that the after-hours price change, $x = f^c_{t_*-1} - p^c_{t-1}$, is an efficient prediction of the expected overnight change in the home exchange price, $y^e = E\{p^o_t - p^c_{t-1} \mid t* - 1\}$, where the superscripts denote opening "o" and closing "c" prices on the relevant exchange. But assume that the cost of a transaction in any exchange is $(\tau/2)$ percent of the unit transaction price, so the cost of a round-trip transaction is τ percent. If profitable cross-exchange investment opportunities do not remain unexploited, the price in the home exchange will move to fully reflect the after-hours price change only when the discrepancy between the expected value of the overnight change of the home exchange price y^e and the observed value of the after-hours foreign price change x exceeds τ. That is, equalization of cross-market valuations is subject to a "no-informative-trading" or "zero-news" interval.

$$|y^e - x| \leq \tau, \tag{6}$$

defined by the round-trip transaction cost τ. For example, if the foreign price is significantly higher than the expected opening home price, investors would continue to sell in the foreign market and anticipate purchases in the home market until the condition of equation (6) is satisfied. However, transactions on the basis of cross-market information would not be profitable if the perceived price discrepancy is less than the cost of transactions τ.[9]

A. Sorting the Size of After-Hours News on the Basis of Transaction Costs

Given data on individual stock prices, one direct way to evaluate the impact of transaction costs on the cross-market transmission of after-hours

[9] In addition to the possibility of liquidity trading, an after-hours (percentage) price change smaller than τ can occur because the information summarized in the price change, $f^c_{t_*-1} - p^c_{t-1}$, is the cumulative "sum" of all information received between the close of the home exchange and the close of the foreign exchange. Also, as noted later, foreign transaction costs may exceed U.S. transaction costs due to the existence of foreign transfer taxes. Consequently, an after-hours foreign price change of zero could be consistent with a latent cumulative price change due to overnight "news" that exceeded U.S. transaction costs but not foreign transaction costs; this case is observationally equivalent to information received after the foreign close and before the U.S. open.

price movements is to sort the after-hours price news into two regressor "bins": after-hours price changes greater than or equal to the transaction cost, $|x| \geq \tau$, and after-hours price changes less than the transaction cost, $|x| < \tau$. If cross-market information is subject to the zero-news interval defined by equation (6), then the home exchange response to regressors in the first bin should be unity, and the response to regressors in the second bin should be zero.[10]

The two-bin sort of after-hours news is based on the following estimates of transaction costs: according to Table 3.7 of the Brady (1988) report, brokerage commissions in 1986 were about 7.5 cents per share for institutional traders or about 0.2 percent, based on an average price per share of $38.60 in 1986 [p. 75 of the *New York Stock Exchange 1988 Fact Book*]; combined with an average bid-ask spread of around 0.4 percent, reported for representative S&P 500 stocks in September 1987 by Bodurtha (1987), estimated round-trip transaction costs for institutional investors in the U.S. were assumed to be in a neighborhood of 0.8 percent; in the London and Tokyo stock markets, additional transaction costs are imposed because of transfer taxes; a seller in Tokyo pays a transfer tax of 0.55 percent, while a buyer in London pays a "stamp duty" of 0.5 percent (see Roll (1988) and Spicer and Oppenheim (1988)); because, on average, an investor may be expected to sell in a given market half of the time, half of these transfer taxes are added to estimated transaction costs in the U.S. market to arrive at an estimate of τ. Thus, estimates of relevant bi-market transaction costs τ are: 1.05 percent for London-New York transactions, 1.075 percent for Tokyo-New York transactions, and 1.325 percent for Tokyo-London transactions.

Two variations of regressions based on transaction cost sorts of after-hours price movements are presented in Table IV. In Panel A of Table IV, close-to-close changes in the New York price of an individual stock are regressed on transaction cost sorts of the after-hours Tokyo price change and the after-hours London price change (defined as the change from the Tokyo close to the London 7:00 A.M. (EST) "close"). In all but one case (the American Express response to London after-hours news), the New York responses are in a statistical neighborhood of unity for large after-hours news and then fall near a statistical neighborhood of zero for small after-hours news.[11] A similar pattern of response behavior is evident in the Dow Jones index regressions presented in Panel B of Table IV. Thus, the results in Table IV are generally consistent with the transaction cost conjecture that the home price response to after-hours news that exceeds transaction costs, $|x| \geq \tau$, should be near unity, and the home price response to after-hours news that is smaller than transaction costs, $|x| < \tau$, should be near zero. A more direct translation of

[10] Use of transaction costs to sort directly the after-hours news of foreign listings was suggested by an anonymous referee.

[11] As noted earlier, the statistical power of individual stock regressions using close-to-close price changes is relatively low. In the case of Table IV, many of the estimated New York responses by individual stocks to small after-hours price changes, $|x| < \tau$, are not statistically different (using 5 percent critical values) from either zero or unity.

Table IV

Transaction Cost Sorting of Responses to After-Hours Price Movements Using Daily Data from October 21, 1987 through June 15, 1988

The table estimates separate responses of home market prices to partitions of after-hours foreign price changes; one partition of the foreign price regressors contains price changes less than ($<$) the estimated round-trip transaction costs τ, and the remaining partition contains foreign price changes that are greater than or equal to ($\geq$) τ. Standard deviations of the coefficients of the partitioned regressors are listed in parentheses. Panel A displays estimated responses of New York close-to-close percentage changes of individual share prices to partitions of the percentage after-hours changes of the corresponding share prices on the Tokyo and London exchanges. Panel B presents, respectively, estimates of the response of the London Dow Jones index to partitions of the after-hours news provided by the Tokyo Dow Jones index and the response of the New York Dow Jones index to partitions of the after-hours news provided by the London Dow Jones index. In the cases of both individual share prices (Panel A) and the Dow Jones indexes (Panel B), ΔTK denotes the percentage deviation of the Tokyo close from the preceding New York close, and ΔLN denotes the percentage deviation of the London "close" at 7:00 A.M. (EST) from the preceding Tokyo close.

		Tokyo Price Changes		London Price Changes				
	Intercept	$\lvert\Delta TK\rvert \geq \tau$	$\lvert\Delta TK\rvert < \tau$	$\lvert\Delta LN\rvert \geq \tau$	$\lvert\Delta LN\rvert < \tau$	N	R^2	DW
Panel A: Individual stock price responses (close-to-close) to after-hours price changes								
American Express	.001	.700	.681	.802	1.148	123	.25	2.28
	(.003)	(.228)	(.609)	(.193)	(.258)			
DuPont	.000	1.316	−.122	1.100	.465	126	.25	1.71
	(.002)	(.220)	(.516)	(.238)	(.413)			
Kodak	−.001	.816	.558	.904	.042	127	.17	2.09
	(.002)	(.207)	(.439)	(.188)	(.231)			
Exxon	.001	1.375	1.125	1.190	.644	104	.20	1.52
	(.002)	(.297)	(.624)	(.295)	(.427)			
General Motors	.001	1.185	.531	.888	.129	123	.14	1.88
	(.002)	(.279)	(.621)	(.247)	(.506)			
IBM	−.002	1.080	.425	.780	−.155	127	.21	1.99
	(.002)	(.200)	(.437)	(.275)	(.283)			
Panel B: Dow Jones index responses to index changes in preceding market								
London index response to Tokyo index	−.001	.751	.212	-	-	94	.33	1.15
	(.001)	(.116)	(.121)					
New York index response to London index	.002	-	-	1.161	.382	94	.37	1.53
	(.001)			(.165)	(.159)			

the transaction cost interpretation to the issue of time-variation in cross-market correlations is presented in the next subsection.

B. A Construction Based on Sample Moments

Several studies, such as Bennett and Kelleher (1988) and King and Wadhwani (1990), have noted that cross-market correlations tend to be positively correlated with measures of price volatility. An intuitive way to explain this association in the context of transaction costs is based on the following characterization of "average" news. If after-hours price changes are approximately normally distributed with a zero mean and if average news is defined by the expected absolute value of the after-hours price change, $E\{\mid x \mid\}$, then average news is proportional to the standard deviation of the after-hours price change σ_x where

$$E\{\mid x \mid\} = (2/\pi)^{1/2}\sigma_x. \tag{7}$$

Combining equation (7) with the effective zero-news interval condition of (6) suggests that the average home exchange response to after-hours news will be closer to zero if the standard deviation of after-hours price changes σ_x is small relative to the average transaction cost τ and closer to unity of σ_x is large relative to τ. Consequently, the size of the average home exchange response to after-hours news depends on the *joint* interaction of the representative *transaction cost* and the *volatility* (standard deviation) of after-hours price movements.

A more formal derivation of the downward "bias" that is expected for the average home exchange response to after-hours news follows. The home response derived under explicit volatility and transaction cost assumptions will be denoted by β_τ to distinguish it from a direct regression estimate of the home exchange response $\hat{\beta}$.

Given the zero-news interval condition of (6), a linear projection of overnight changes in the home price y on after-hours foreign price changes x will be proportional to the integral of the density of x that excludes the zero-news interval, $\mid x \mid < \tau$. The variance of x and the covariance of x and y can be expressed as:

$$\begin{aligned}\operatorname{var}(x) &= E(x^2)\\ &= \Pr(\mid x \mid \geq \tau) \cdot E(x^2\colon \mid x \mid \geq \tau) + \Pr(\mid x \mid < \tau) \cdot E(x^2\colon \mid x \mid < \tau)\end{aligned} \tag{8.a}$$

$$\operatorname{cov}(y, x) = \Pr(\mid x \mid \geq \tau) \cdot E(x^2\colon \mid x \mid \geq \tau) + \Pr(\mid x \mid < \tau) \cdot 0. \tag{8.b}$$

Thus, a transaction cost interpretation of a regression of overnight home exchange price changes y on after-hours foreign price changes x is:

$$\begin{aligned}\beta_\tau &= \operatorname{cov}(y, x)\operatorname{var}^{-1}(x)\\ &= \Pr(\mid x \mid \geq \tau) \cdot \frac{E(x^2\colon \mid x \mid \geq \tau)}{E(x^2)}\\ &= 1 - \Pr(\mid x \mid < \tau) \cdot \frac{E(x^2\colon \mid x \mid < \tau)}{E(x^2)}.^{12}\end{aligned} \tag{9}$$

[12] It can be shown that β_τ has the expected boundary properties:

$$\lim_{\tau \to \infty} \beta_\tau = \lim_{E(x^2) \to 0} \beta_\tau = 0,$$

$$\lim_{\tau \to 0} \beta_\tau = \lim_{E(x^2) \to \infty} \beta_\tau = 1.$$

Under the simplifying assumption that after-hours percentage price changes are approximately normally distributed,[13] the transaction cost response β_τ may be reformulated as

$$0 \leq \beta_\tau = 2 - 2\left[\Phi(\tau/\sigma_x) - \mathrm{f}(\tau/\sigma_x)\tau/\sigma_x\right] \leq 1, \tag{10}$$

where $\Phi(\cdot)$ denotes the unit normal cumulative distribution function, $\mathrm{f}(\cdot)$ is the unit normal probability density function, σ_x is the standard deviation of (percentage) after-hours foreign price changes, and we have employed a result in Johnson and Kotz (1970) for the variance of a truncated normal distribution (p. 83), viz.

$$\frac{E(z^2: |z| < k)}{E(z^2)} = 1 - \frac{2\mathrm{f}(k)k}{2\Phi(k) - 1}. \tag{11}$$

As summarized in equation (10), transaction costs provide a straightforward interpretation of time-varying estimates of β_τ. Under normality, it can be shown that β_τ is monotone increasing in $E\{x^2\}$ and monotone decreasing in τ. Consequently, *for constant transaction costs τ the estimated predictive efficiency of after-hours foreign pricing will appear to vary directly with observed foreign price volatility*, represented under the normal distribution assumption by the standard deviation of after-hours price changes σ_x.

One method of assessing the empirical validity of this formulation of the transaction cost interpretation is to contrast subperiod constructions of β_τ, based on subperiod estimates of σ_x and τ, with the subperiod regression estimates $\hat{\beta}$ that were presented in Table I. Estimates of the standard deviations σ_x of foreign (and New York) percentage price changes are presented in Table V. The standard deviations of after-hours price changes were exceptionally high immediately following the 1987 crash. Although the price volatility of the Tokyo index apparently fell to near-normal levels by the second 10-day interval after the crash, the price volatility of the London index subsided more gradually.

Approximate subperiod estimates of the transaction cost variant of the New York exchange response β_τ are constructed by substituting the earlier estimates of bi-market transaction costs into equation (10), along with relevant subperiod estimates of the standard deviations of after-hours Tokyo and London price changes denoted by σ_{TK}^{ah} and σ_{LN}^{ah}, respectively, in Table V. Also, lower and upper boundaries of 95 percent confidence intervals for the esti-

[13] There is a considerable literature on distributions that more adequately capture the leptokurtic characteristic of auction prices, such as the classic work by Mandelbrot (1963) on stable Paretian distributions. However, this refinement is less important in the present context since equation (9) is directed at integration of the density in a small τ-neighborhood of the median price change. A contrast of standardized normal and Cauchy distributions in Johnson and Kotz (1970) suggests that differences in the cumulative distributions are negligible for τ/σ_x less than unity.

Table V

Standard Deviations of Overnight, Trading Day, and After-Hours Percentage Changes in Stock Prices Using Daily Data from October 21, 1987 through June 15, 1988

The table presents sample partitions of the standard deviations of trading day (open-to-close) and overnight (close-to-open) percentage changes of the New York Dow Jones index and after-hours (New York close-to-foreign close) percentage changes of Dow Jones indices constructed by the authors from available shares of Dow Jones stocks listed on the Tokyo and London exchanges. The trading days in the sample partitions of the New York prices are matched to the available trading days incorporated in the sample partitions of the Tokyo (Panel A) and London (Panel B) data. The standard deviation of Tokyo after-hours price changes is denoted by σ_{TK}^{ah}; the standard deviation of London after-hours price changes by σ_{LN}^{ah}; and the standard deviations of New York overnight and trading-day price changes by σ_{NY}^{on} and σ_{NY}^{td}.

	Trading Days Following October 20, 1987					
	Full Sample	Days 1–30	Days 30+	Days 1–10	Days 11–20	Days 20+
	Panel A: Data used for Tokyo regressions					
σ_{TK}^{ah}	.007	.013	.005	.022	.004	.005
σ_{NY}^{on}	.013	.029	.006	.041	.011	.008
σ_{NY}^{td}	.016	.021	.014	.033	.017	.014
	Panel B: Data used for London regressions					
σ_{LN}^{ah}	.007	.014	.004	.021	.013	.004
σ_{NY}^{on}	.011	.024	.006	.040	.018	.007
σ_{NY}^{td}	.014	.019	.013	.032	.015	.013

mated transaction cost response β_τ are provided under the assumption that the after-hours price changes are normally distributed, and, thus, the sample variance estimates are χ^2 distributed.[14]

As reported in Panel A of Table VI, the time sequences of New York responses to after-hours price information generated by the transaction cost interpretation β_τ are generally statistically consistent with the time-varying profiles of New York responses estimated by direct regression $\hat{\beta}$ in Table I. In the case of New York responses to Tokyo prices, all of the responses predicted by the transaction cost estimates β_τ are contained by the 95 percent confidence intervals of the regression estimates $\hat{\beta}$. Reciprocally, most of the regression estimates $\hat{\beta}$ are contained by the 95 percent confidence intervals of the transaction cost estimates β_τ. In the case of estimated New York responses to after-hours changes in London prices, both the regression $\hat{\beta}$ and

[14] Since β_τ is monotone increasing in σ_x, confidence intervals for β_τ were constructed by applying equation (10) to the confidence interval boundaries for σ_x^2. The confidence intervals are conservative estimates of the true uncertainty since no allowance was made for errors in measuring transaction costs.

Table VI

Contrasts of Regression $\hat{\beta}$ and Transaction Cost β_τ Responses of New York and Foreign Prices

Regression estimates $\hat{\beta}$ of responses of home share prices to after-hours foreign price movements are contrasted with responses predicted by transaction cost constructions β_τ. The latter assume that the home market efficiently utilizes after-hours information but does not respond to after-hours price changes that are less than the estimated round-trip transaction cost τ. The distribution of effective after-hours news (which is truncated in a $\pm\tau$ region about zero) is based on the standard deviations of price changes presented in Table V. Lower and upper boundaries of the 95 percent confidence intervals of both the regression estimates and the transaction cost constructions are reported in parentheses. Panel A displays the New York overnight (close-to-open) percentage responses of the Dow Jones index to after-hours (New York close-to-foreign close) percentage changes of Dow Jones indices constructed for the Tokyo and London exchanges. The foreign Dow Jones indices are based on available Dow Jones shares listed on the foreign exchanges. Panel B presents estimates of the **reverse** responses of the foreign Dow Jones indices (New York open-to-foreign close) to trading day (open-to-close) changes of the U.S. Dow Jones index.

	Trading Days Following October 20, 1987				
	Days 1–30	Days 30+	Days 1–10	Days 11–20	Days 20+
Panel A: Responses of New York overnight pricing to foreign after-hours price changes					
New York Responses to Tokyo Prices					
$\hat{\beta}$	.934	.041	.912	1.872	.094
	(.524; 1.344)	(−.418; .500)	(.487; 1.337)	(−.945; 4.689)	(−.355; .543)
β_τ	.877	.202	.971	.065	.202
	(.517; .983)	(.044; .471)	(.637; .999)	(.000; .958)	(.053; .461)
New York Responses to London Prices					
$\hat{\beta}$	1.132	.478	1.633	.939	.408
	(.887; 1.377)	(−.070; .886)	(1.276; 1.990)	(.582; 1.296)	(.083; .733)
β_τ	.905	.075	.969	.884	.075
	(.631; .985)	(.011; .264)	(.462; 1.000)	(.297; .995)	(.020; .318)
Panel B: Responses of foreign pricing to New York trading day price changes					
Tokyo Responses to New York Prices					
$\hat{\beta}$	.805	.911	.715	1.077	.918
	(.653; .957)	(.809; 1.013)	(.540; .890)	(.665; 1.488)	(.818; 1.018)
β_τ	.991	.940	.967	.899	.899
	(.829; 1.000)	(.249; .999)	(.825; .996)	(.792; .956)	(.805; .955)
London Responses to New York Prices					
$\hat{\beta}$	1.021	1.119	.857	1.147	1.135
	(.869; 1.173)	(1.027; 1.211)	(.661; 1.053)	(.831; 1.463)	(1.047; 1.223)
β_τ	.991	.921	.959	.884	.884
	(.775; 1.000)	(.380; .997)	(.811; .944)	(.784; .944)	(.814; .953)

transaction cost β_τ estimates of the response to after-hours London price changes remain near unity throughout the first 30 trading days following the crash before dropping to a statistical neighborhood of a zero response. Although there is somewhat less correspondence between the transaction cost and regression point estimates for the London data, the confidence intervals generally overlap and reveal the same pattern of a statistically significant decline in New York responses to after-hours price movements.[15]

C. The Persistent Transmission of U.S. News

Further evidence on the transaction cost interpretation may be obtained by examining the *reverse* flow of information, *from* daily New York price changes *to* succeeding prices in the Tokyo and London markets, to see if the relevant transaction cost β_τ and regression $\hat{\beta}$ estimates agree on after-hours responses to New York news.

Panel B of Table VI presents estimates of the Tokyo and London responses to New York trading day price changes, along with boundaries of the relevant 95 percent confidence intervals. Here, the dependent variables of the regressions to estimate foreign market responses $\hat{\beta}$ are the (percentage) differences between the foreign close and the prior New York open, and the regressors representing New York news are provided by open-to-close (percentage) changes in New York prices. In contrast to the time-varying pattern of New York responses to foreign price news, estimates in Panel B of Table VI suggest that foreign responses to New York news generally remained in a statistical neighborhood of unity throughout the sample.

Both the regression $\hat{\beta}$ and transaction cost β_τ estimates of foreign responses are consistent with the persistent efficiency of cross-market transmissions from New York trading day price changes. The foreign responses estimated by the transaction cost method β_τ are: slightly above the regression 95 percent confidence intervals estimated for the Tokyo response in the first 30 days and in the first 10 days; below the regression 95 percent confidence intervals estimated for the London response beginning 20 or 30 days after the October crash; and otherwise contained by the estimated regression confidence intervals.

IV. Variations in the Global Distribution of News

The empirical results in this paper indicate that in the immediate aftermath of the October 1987 stock market collapse, prices of U.S. stocks with listings on the Tokyo and London exchanges were tightly linked across international capital markets. In particular, after-hours price movements were reflected in home prices in all exchanges. This is the expected result if international

[15] As noted earlier, the extraordinarily large response estimated by the regression method for the first 10-day interval $\hat{\beta} = 1.633$ suggests a hypersensitive response by New York investors to after-hours London price changes. The degree of hypersensitivity may be understated by the spread between the regression and transaction cost responses because the latter β_τ was constructed on the assumption of a constant transaction cost τ throughout the sample. Compilations reported in Bodurtha (1987) suggest that bid-ask spreads may have increased by 50 percent to 100 percent in the initial aftermath of October 19. Comparable increases in "touches" (spreads between the highest bid and lowest offer) were observed in the London exchange in the immediate wake of the crash (see Bank of England (1988)).

capital markets are efficient, or not segmented. This behavior was not observed, however, in stock price data collected for trading days further removed from the collapse. In intervals of more typical price volatility, transmission from after-hours foreign price movements to opening New York prices was significantly weaker or entirely absent, whereas foreign prices continued to fully reflect the New York trading day news.[16] Both the asymmetry of cross-market responses and the time variation in New York responses appear to be due to the impact of transaction costs on the profitability of responding to after-hours news. Transaction costs provide a barrier to international transmissions of small (but not large) price signals.

Interpretation of the asymmetry of market transmissions in normal trading periods is based on the uneven global distribution of relevant news. In less turbulent intervals, such as the sample subperiod beginning 30 trading days after the crash, the preponderance of economic news regarding U.S. stocks is publicly released during New York trading hours. This uneven clustering of news relevant for U.S. stocks is illustrated in the third column of Table V where the volatility of New York trading day prices is generally about three times larger than the volatility of foreign after-hours price changes and about twice the overnight volatility.[17]

Interpretation of the time variation in New York responses is based on changes in the global distribution of news as illustrated in the historical snapshots of Table V. In the case of the first 10 trading days following the October 19, 1987 crash, the ratio of after-hours price volatility to New York trading day price volatility nearly doubled, as shown in the fourth column of Table V. Since it may be presumed that the dominant "news" in the days immediately following the crash was the volatile behavior of other investors, it is not surprising that the sources of effective investor information immediately after the October crash did not conform to the informational rhythm of a more representative international trading day.[18] That is, relevant information about the behavior of other investors was more globally dispersed

[16] This asymmetry in share price comovements is similar to that demonstrated by Eun and Shim (1989) who estimate a multivariate VAR of daily price movements (in local currency units) of nine international exchange indices over a pre-crash sample, January 1980 through December 1985. With regard to the news contained in share price movements, the authors conclude, "Innovations in the U.S. stock market are rapidly transmitted to other markets in a clearly recognizable pattern, whereas no single foreign market can significantly explain the U.S. market movements" (p. 243).

[17] These sample estimates of the ratios of daytime volatility to overnight or after-hours volatility during periods of "normal" price volatility are roughly consistent with estimates reported in the Stoll and Whaley (1990) study of 1982–1986 transactions on the New York Stock Exchange. Using the standard deviation of returns as a measure of price volatility, results in Table VI of Stoll and Whaley (1990) indicate that the average ratio of daytime volatility to overnight volatility was about 2.3 for all stocks and about 2.7 for stocks in the quintile with the largest trading volume.

[18] The importance of investors' views of the attitudes of other investors is indicated in post-crash surveys of participants in U.S. markets by Shiller (1987) and in Japanese markets by Shiller, Konya, and Tsutsui (1988).

because the activities of investors were not (and are not) confined by the geographical boundaries of trading hours of home exchanges.[19]

Finally, this paper has provided some evidence that investor transaction costs appear to have sizeable impacts on the asymmetry and episodic variations of cross-market correlations. Other testable implications of investor transaction costs may warrant scrutiny. As discussed above, transaction costs tend to "hollow out" the frequency distribution of investor short-run responses to accumulated news; thus, changes in transaction costs may alter the kurtosis or relative frequency of large movements in share prices. Also, if global deregulation of financial markets in the 1980's induced a secular decline in effective transaction costs then, *ceteris paribus*, cross-market correlations should have increased in the last decade.[20]

[19] An increase in time-zone trading by foreign investors during the October crash is noted in Bank of England (1988). Models of strategic intraday trading patterns are explored by Admati and Pfleiderer (1988) and of strategic time-zone trading in dual-listed stocks by Freedman (1989).

[20] Evidence of moderate strengthening in cross-market correlations during the 1980's is presented in Bennett and Kelleher (1988) and Kupiec (1990).

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