



American Finance Association

Primes and Scores: An Essay on Market Imperfections

Author(s): Robert A. Jarrow and Maureen O'Hara

Source: *The Journal of Finance*, Vol. 44, No. 5 (Dec., 1989), pp. 1263-1287

Published by: Blackwell Publishing for the American Finance Association

Stable URL: <http://www.jstor.org/stable/2328642>

Accessed: 26/11/2009 07:31

Your use of the JSTOR archive indicates your acceptance of JSTOR's Terms and Conditions of Use, available at <http://www.jstor.org/page/info/about/policies/terms.jsp>. JSTOR's Terms and Conditions of Use provides, in part, that unless you have obtained prior permission, you may not download an entire issue of a journal or multiple copies of articles, and you may use content in the JSTOR archive only for your personal, non-commercial use.

Please contact the publisher regarding any further use of this work. Publisher contact information may be obtained at <http://www.jstor.org/action/showPublisher?publisherCode=black>.

Each copy of any part of a JSTOR transmission must contain the same copyright notice that appears on the screen or printed page of such transmission.

JSTOR is a not-for-profit service that helps scholars, researchers, and students discover, use, and build upon a wide range of content in a trusted digital archive. We use information technology and tools to increase productivity and facilitate new forms of scholarship. For more information about JSTOR, please contact support@jstor.org.



Blackwell Publishing and American Finance Association are collaborating with JSTOR to digitize, preserve and extend access to *The Journal of Finance*.

<http://www.jstor.org>

Primes and Scores: An Essay on Market Imperfections

ROBERT A. JARROW and MAUREEN O'HARA*

ABSTRACT

This paper investigates the reported relative mispricing of primes and scores to the underlying stock. Given transaction costs, we establish arbitrage-based bounds on prime and score prices. We then develop a new nonparametric statistical technique to test whether prime and score prices violate these bounds. We find that prime and score prices do exceed stock prices, and often by a considerable amount. We demonstrate that this increased value is most likely due to the score's ability to save on the costs of dynamic hedging. We also show how short sale and trust size constraints impede the ability to arbitrage price disparities.

IN AN EFFICIENT MARKET the price of a security reflects the value of its underlying assets. It should not be possible, therefore, to create value simply by splitting a security into different parts. Since the underlying assets are unchanged, the packaging of the cash flows should not matter. The recent introduction of primes and scores suggests that this is not always the case. By separating a security's cash flows into a dividend-based component (the prime) and an appreciation or option-based component (the score), developers of these securities have seemingly created two assets whose value separately exceeds that of the original stock. Since the prime and the score are based on the same underlying asset, this raises a puzzling question about the actual efficiency of securities markets.

This paper investigates this anomaly by examining whether primes and scores are actually "mispriced" and, if so, why. Given the costs of transacting in the prime, score, and stock, we establish bounds on prime and score prices that are consistent with no arbitrage opportunities. We then develop a new nonparametric statistical technique to test whether prime and score prices violate these bounds. This nonparametric procedure is designed both to confront the measurement error problems plaguing the literature (i.e., nonobservable bid/ask prices and nonsimultaneous prices; see Phillips and Smith (1980), Bhattacharya (1983), and Bookstaber (1981)) and to avoid the estimation of round trip transaction costs (commissions).

We find the surprising result that prime and score prices exceed the price of the underlying stock, often by a considerable amount. Using time-series data, we demonstrate that significant price differentials exist for each firm in our sample.

* S. C. Johnson Graduate School of Management, Cornell University. We would like to thank Robin Brenner for computer assistance and David Easley, B. Jesper Christensen, James MacBeth, Dale Morse, René Stulz, Robert Whaley, an anonymous referee, and seminar participants at Cornell, Michigan, Ohio State, and the Western Finance Meetings (June 1988) for helpful comments. Financial support from the Mobil Oil Foundation is gratefully acknowledged.

We then investigate whether these price differentials are consistent with the existence of arbitrage opportunities. We show that, while transaction costs generally preclude such opportunities for the average trader, this is not true for traders subject to lower costs. One limit to any arbitrage profits, however, is the small daily volume of trading in primes and scores. We also show how short sale constraints and limits on the size of the trust may restrict the ability to arbitrage, leaving prime and score prices free to deviate from the underlying stock price.

One question suggested by these results is why primes and scores should be overpriced in the first place. We examine three alternative explanations for this price behavior. Specifically, we investigate whether a market completeness, a transactions cost, or a tax-based argument can explain why primes and scores should be more valuable than the underlying stock. Our research identifies the transaction costs of dynamic hedging as a possible reason for overpricing. We argue that the long-term option created by the score avoids the cost of replicating such an option via dynamic hedging, causing the score to be more valuable than predicted by a standard option pricing model. Consequently, this ability to save on transaction costs may make the prime plus score combination more valuable than the underlying stock.

One way to interpret our results is that the behavior of prime and score prices is a product of several market imperfections. By focusing on these market imperfections, our work demonstrates why value can arise from splitting securities and how a price differential can persist. The development of other dichotomized securities suggests that similar imperfections may exist in other markets. For example, the separation of mortgage-backed securities (MBS) into interest-only (IO) and principal-only (PO) components provides a debt-based analogy to the equity instruments analyzed here. Since the IO and PO typically sell at a premium to the underlying MBS, the approach developed in this paper may prove useful in understanding the behavior of these debt instruments. More recently, the proposal to further divide equity into unbundled stock units suggests that such segmentation may be an important characteristic of future securities markets.

An outline of this paper is as follows. Section I describes primes and scores and derives the arbitrage-based price restrictions. Section II presents the statistical methodology for testing these pricing relationships. Section III tests these relationships using daily closing price observations from the *Wall Street Journal*. Section IV applies option pricing theory to value the score and to investigate various hypotheses concerning the price differences. Finally, Section V concludes the paper.

I. The Structure of Primes and Scores

Primes and scores are not originally issued securities but are created through establishment of a trust. The trust accepts shares of common stock in a specified company and issues in exchange a unit of the trust. The trust has a maturity of five years, at which time the outstanding units are reconverted to the underlying stock. Each unit contains a prime and a score. The prime component receives all

dividend payments and any increase in the stock price up to a termination value. The score receives any appreciation above the termination value. The termination price is set at the beginning of the trust and has generally been at a 20–25% premium to the current stock price. A trust may not accept further shares once the stock price exceeds the termination value.

The two-tiered structure of the trusts is reminiscent of a dual purpose fund. One difference is that the trust will exchange a prime and a score for the underlying stock at any time during trading hours at no cost. This redemption feature allows holders to receive at least the value of the stock at any time.¹ A second difference is that holders of the trusts can trade separately each component. The prime, score, and unit for each trust are separately listed on the American Stock Exchange (ASE).

The first trust was established in 1983 by Americus Shareowners Service Corporation for the prevestiture AT&T shares. Americus Trusts for Exxon, DuPont, American Home Products, Merck, and Bristol-Myers followed in 1985 and 1986. There are currently 27 Americus Trusts, although most have only been trading since August 1987. Because of a recent IRS ruling on the taxation of the trusts, no future trusts are expected to be created.² Each existing trust is limited to holding 5% of a corporation's outstanding stock. Once a trust reaches that level it is closed, and no further primes and scores in that company can be created.

The costs of the trust fall into two categories. There is an initial deposit fee per share tendered to the trust, with per share charges decreasing with number of shares. There is also an annual fee of \$0.06 per unit assessed on the prime component (except Exxon, which is \$0.05 per unit). There is no fee for exchanging units of the trust for the underlying stock.

To focus on the behavior of prime and score prices, we incorporate these features into a simple model. Let the trust have a maturity of T and a termination price of K . We define S_t to be stock price, P_t prime price, and C_t score price, all at time t . Since the prime receives all dividends, the score can be viewed as a long-term European call option on S_t with exercise price K and exercise date T . At maturity, therefore, the score price is $C_T = \max(S_T - K, 0)$ and the prime price is $P_T = \min(S_T, K)$.

Because the prime and the score have the same composite cash flows as the stock, there are bounds on how widely their prices can diverge. If the combined prime and score price differs from the stock's price, then arbitrageurs will buy

¹ This redemption policy is clearly stated in the Prospectus for each Trust: "On any day on which the NYSE is open for trading prior to the termination of the Trust, any unseparated unit may be redeemed, and any Prime and Score component may be recombined to form a unit which may be redeemed, for an in kind distribution equal to the Net Asset Value Per Unit. No fee will be charged for the redemption of any Unit. It should be noted, however, that holders of separate components may have to pay customary brokerage fees and commissions to acquire complementary components to recombine Units."

² Specifically, on May 2, 1984, the IRS published proposed amendments to Treasury Regulation Section 301.7701-4(c) that classify the Trust as an association taxable as a corporation for federal income tax purposes. These proposed amendments were made final on March 21, 1986. Because this ruling results in double taxation, it is expected that no new trusts will be formed.

the cheaper and sell the more expensive security(ies), providing a risk-free profit. In the absence of transaction costs, these prices would obviously have to be the same. The presence of transaction costs means that there will be a range of prices that preclude arbitrage opportunities.

To determine the lower bound of prices, suppose the stock is selling at a premium to the prime and the score. An arbitrageur should buy the prime and score, convert them to the stock, and then sell the stock. Because this transaction will be executed immediately, there are no annual management fees to consider. Moreover, since taxes affect each transaction similarly, tax effects need not be explicitly incorporated. The arbitrageur pays trading commissions of c percent on the purchases and sale, but there is no fee for exchanging the prime and score into the stock.

Suppose, instead, that the stock is selling at a discount to the prime and score. There are two potential methods to arbitrage the difference. One approach is to buy the stock, exchange it to the trust for a prime and a score, and then sell the prime and score. The arbitrageur again pays trading commissions on both purchase and sale and also incurs a sales charge of m percent for the creation of the prime and score. To avoid arbitrage profits, then, prices must satisfy

$$-c(P_t + C_t + S_t) \leq P_t + C_t - S_t \leq c(P_t + C_t + S_t) + mS_t, \quad (1)$$

where c reflects both brokerage commissions and the effect of the bid-ask spread.

Because of size limitations on the trust, however, it may not be possible to create additional primes and scores. If the stock is undervalued relative to the prime and score, a second arbitrage approach is to sell the prime and score short, buy the stock, and hold this position until the maturity date at time T . At the maturity of the trust the two sides must be equal, so this allows the arbitrageur to lock in the differential. This strategy incurs both brokerage commissions and the costs of short sale restrictions. The most significant of these is the regulation that proceeds from a short sale be held in a non-interest-bearing account (for the life of the short sale).³ We denote by i the present value of the interest lost as a percent of the short sale. The price bounds consistent with this strategy are then

$$-c(P_t + C_t + S_t) \leq P_t + C_t - S_t \leq c(P_t + C_t + S_t) + i(P_t + C_t). \quad (2)$$

Equations (1) and (2) dictate the interval in which prices can exist without generating arbitrage opportunities. Deviations from these bounds, therefore, would indicate a violation in market efficiency. These transaction cost bands implicitly assume that the arbitrageur acts as a price taker, meaning that he or she can execute a trade at the quoted prices and that his or her trade will not change the price. This subtle aspect of the pricing theory will be relevant to our subsequent results. In the next section we use these bounds to test for inefficiencies in prime and score pricing.

³ The proceeds constraint is part of the Federal Reserve Board's margin requirement rules. Specifically, Regulation T specifies that the proceeds of a short sale must be deposited in a non-interest-bearing account as well as an additional 50% of the short sale amount which may be held in interest-bearing securities. See Cox and Rubinstein (1985) for more details.

II. Statistical Methodology

The price bounds derived above provide a framework for testing for the existence of security mispricing. Unfortunately, the estimation of these bounds may be subject to severe measurement error problems. These measurement errors arise from nonsynchronous price observations, the lack of bid/ask prices, difficulties in determining marginal trading commissions, and even typographical errors in data reporting. (See Phillips and Smith (1980) and Bookstaber (1981) for relevant discussions.) The nonsynchronous data problem is particularly likely to be significant for primes and scores because, as we show later, the trading volume is often quite small.

The procedure often used to overcome these problems is to refine the data set by obtaining transaction prices (see Bhattacharya (1983)) and to be exhaustive in providing differing traders' transaction costs. (See Foster and Oldfield (1987).) For most applications these refined data are difficult to acquire. Detailing actual transaction costs is even more problematic. It may be impossible to determine either the types of traders transacting in a particular market or the precise level of costs each faces. Also, even if these problems were solved, the measurement errors arising from other sources would still remain a problem.

One contribution of this paper is a new nonparametric statistical methodology designed to confront these problems. We develop a procedure that uses part of the data to estimate implicitly the pricing discrepancies arising from noise and trading costs. What motivates this procedure is that the overall distribution of pricing errors can be described as a composite of two underlying distributions, one symmetric with respect to zero and the other asymmetric. Specifically, the data set we test consists of *observed* daily closing prices, which can be decomposed into actual prices and an observation error.⁴ From our pricing equations, the observed price difference, defined as $\hat{\delta}_t$, is bounded by

$$-c(P_t + C_t + S_t) + \epsilon_t \leq \hat{\delta}_t \leq c(P_t + C_t + S_t) + \epsilon_t + M_t, \quad (3)$$

where M_t represents either the short sale cost ($i(P_t + C_t)$) or the cost of arbitraging through the trust (mS_t), and ϵ_t is the observation error.

Arbitraging price discrepancies between the prime and score and the stock involves transaction costs plus noise to both the left and right of zero. However, these are asymmetric since the right-hand distribution also involves the additional costs of either short selling or creating new securities. Since the trust always stands ready to exchange pairs of the prime and score for the stock at *no* cost, any deviation to the left of zero is likely to result from bid/ask spreads, commissions, or nonsimultaneous prices rather than being an arbitrage opportunity. We use the left-hand distribution to estimate these measurement errors and transaction costs and then use these estimates to test whether arbitrage opportunities are present in the right-hand distribution.

Figure 1 illustrates this approach. The figure plots a typical distribution of the

⁴ Although we derive our testing procedure for absolute differences, the same analysis applies to normalized percent differences. The only necessary change is that the subsequent statistical structure is applied to ϵ_t^* and δ_t/S_t rather than ϵ_t and δ_t , with $\epsilon_t^* \equiv \hat{\delta}_t/\hat{S}_t - \delta_t/S_t$.

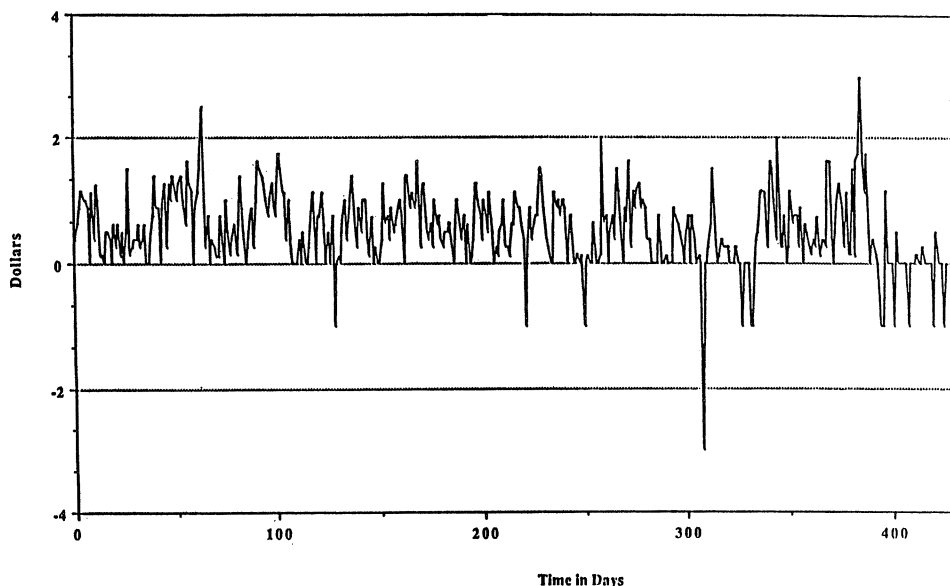


Figure 1. A typical distribution for the daily observed price differences of the prime plus score less stock prices. The “true” prime price, score price, and stock price at time t are denoted P_t , C_t , and S_t , respectively. The percent transaction costs are captured by the fraction c . The symbols $\bar{x}_t$, $\bar{x}_t$ represent those random differences due to transactions costs (bid/ask spreads and commissions) to the left and right of zero, respectively. The ϵ_t is the observation error. The y_t represents those deviations due to short interest costs or trust fees. The β is chosen so that $\beta < \theta_t$, where θ_t is a bound on the support of the observation error. The shaded area in the right tail is the mass remaining beyond the level of the maximum error θ_t , transaction cost adjustment c , and additional costs M_t .

observed price differences between the prime and score and the stock. We define the random variables $\bar{x}_t$ to represent random differences to the left of zero due to transaction costs (bid/ask spreads and commissions) and $\bar{x}_t$ to represent random differences to the right of zero due to these *same* transaction costs. (Formal definitions are given in the Appendix.)⁵ Let ϵ_t be the error term and let y_t represent those deviations to the right of zero due to the additional transaction costs, either the loss of short interest or the management fee. (See equations (2) and (3).) As Figure 1 illustrates, the $x_t + \epsilon_t$ distributions are symmetric except for scale. We use the information conveyed by the left side of the $x_t + \epsilon_t$ distribution to estimate the implied level of transaction costs and noise in the data. We then estimate the level of short interest or management costs (the M_t), adjust for the implied transaction cost and noise, and test to see whether observations remain in the right-hand distribution that are not explained by these factors.

In the Appendix we derive a nonparametric statistical procedure to implement this test. Given assumptions on boundedness, independence, and symmetry (see Appendix), the following lemma holds.

⁵ The exact form of the transaction costs (i.e., proportionate, fixed) is irrelevant for our analysis.

LEMMA 1 (Null Hypothesis): Let $\hat{\delta}_t$ be the observed price difference and let θ_t be some constant bound on the error term. Then

$$\text{prob}(\hat{\delta}_t > \beta + M_t) \leq \text{prob}(\hat{\delta}_t \leq -\beta) \quad \text{for all } \beta \geq \theta_t \geq 0.$$

Proof: See Appendix.

This lemma provides the benchmark for testing when prices deviate beyond the level consistent with transaction costs and measurement errors. To develop a hypothesis test, we first suppose that $\hat{\delta}_t$ for $t = 1, \dots, n$ represents a random sample of n independent, identically distributed observations. This implies that the t subscript on both M_t and θ_t can be omitted. Let $F(\alpha) = \text{prob}(\hat{\delta}_t \leq \alpha)$ be the underlying probability distribution. The null and alternative hypotheses, re-expressed in terms of $F(\cdot)$, are

$$H_0: 1 - F(\beta + M) \leq F(-\beta),$$

$$H_a: 1 - F(\beta + M) > F(-\beta).$$

A consistent estimator for the underlying distribution, $F(\alpha)$, is $F_n(\alpha)$, the sample cumulative distribution. Our hypothesis testing procedure uses this estimator to find a $\hat{\beta}$ such that $F_n(-\hat{\beta}) = 0.05$. This percentage is chosen to ensure that $\hat{\beta} \geq \theta$ with probability one. We then calculate the mass in the right tail, $[1 - F_n(\hat{\beta} + M)]$. If $[1 - F_n(\hat{\beta} + M)]$ is much larger than 5% (i.e., $F_n(-\hat{\beta})$), we reject the null hypothesis that the price difference is due to standard transaction costs and noise. Otherwise, it is accepted.

To complete the analysis, we derive the sampling distribution for the estimators to obtain the appropriate confidence regions. By using the sample distribution for $F_n(\alpha)$ for fixed α , a binomial distribution, the following lemma is proven in the Appendix.

LEMMA 2 (Confidence Regions): Given $\hat{\beta} \geq \theta$ such that $F_n(-\hat{\beta}) = 0.05$, for large n (say greater than 50), the $(1 - \alpha)$ percent confidence interval for $1 - F_n(\hat{\beta} + M)$ lies within $[0, 1 - L/n]$, where

$$L \equiv -Z_{\alpha/2}[n(1 - q)q]^{1/2} + (1 - q)n,$$

$$q \equiv ([Z_{\alpha/2}^2 + 0.1n] + Z_{\alpha/2}[Z_{\alpha/2}^2 + 0.19n]^{1/2})/2[n + Z_{\alpha/2}^2], \quad \alpha \in (0, 1),$$

and $Z_{\alpha/2}$ is defined by $\Phi(Z_{\alpha/2}) = 1 - \alpha/2$ for Φ the cumulative normal distribution.

Proof: See Appendix.

Lemma 2 can be explained as follows. Under the null hypothesis, we will observe more than $(1 - L/n)$ percent of the sample distribution to the right of $\hat{\beta} + M$ with probability less than α . Hence, if more than $(1 - L/n)$ percent of the sample distribution lies to the right of $\hat{\beta} + M$, we reject the null hypothesis with at least a $(1 - \alpha)$ percent confidence level. These confidence intervals are conservative, in the sense that the true confidence level is at least $(1 - \alpha)$ percent and probably much greater. Consequently, we report only the 90% confidence levels ($\alpha = 0.1$) in the next section.

Several features of this testing methodology deserve comment. First, we incorporate transaction costs in our testing procedure without having to estimate directly their magnitude. This allows our testing procedure to include the simultaneous effects of the bid/ask spread and commissions that induce bias in the existing procedures. (See Phillips and Smith (1980).) Second, the nonsimultaneity of the observed prices is also incorporated within the procedure. The neglect of this noise has potentially mis-specified existing approaches. (See Bookstaber (1981).) Third, the procedure is nonparametric. It does not require normally distributed errors, yet suitable confidence regions can be derived. This is important since the deviations are partially bound by arbitrage considerations (see equations (1) and (2)), so normality does not apply. Finally, our testing procedure is designed to minimize the probability of finding significance when it does not, in fact, exist. This biases our results, however, in favor of market efficiency and against finding arbitrage opportunities. Our results in the next section should be interpreted with this conservative bias in mind.

Although we apply this technique to primes and scores, the methodology is usable whenever the data satisfy the boundedness and symmetry conditions given in the Appendix. Potential applications include ex-dividend stock price behavior, option pricing put-call parity relationships, IO and PO pricing, and futures-spot arbitrage restrictions. We illustrate these conditions in terms of absolute differences above, but, in some circumstances (perhaps here), the structure is more likely to be satisfied by proportional differences. The same analysis applies with the appropriate modifications.

III. The Empirical Behavior of Prime and Score Prices

Because primes and scores are new, the data available for testing are limited. In particular, only 27 companies have primes and scores, with most of these trading only since August 1987. To avoid the difficulties of an unduly small number of observations, we restrict our analysis to those primes and scores listed before June 1987. Our sample consists of daily stock closing price observations from the *Wall Street Journal* for the prime, score, and stock on Exxon, Bristol-Myers (B-M), DuPont, Merck, and American Home Products (AHP).⁶ The sample period is from the initiation of the particular trust to June 1987.

To test for pricing discrepancies, we first calculated the differences between the daily stock price and the daily combined prime and score prices. A typical plot of these for one company is given in Figure 2. As expected, the differences exhibit variability, reflecting noise, and other factors noted in the previous section.⁷

⁶ We did not include AT&T in our sample for two reasons. First, there was some concern that uncertainty over the divestiture would introduce biases into the data. A more important concern is that, because AT&T primes and scores were the first such instruments, a start-up bias could be present. To avoid this problem, we used those primes and scores introduced in 1985 and 1986.

⁷ Figure 2 suggests that the differences satisfy the maintained i.i.d. hypothesis. It is well known that closing prices may exhibit serial correlation due to bid/ask spreads and rounding errors; see Ball (1988). Taking *differences* of related closing prices, however, could neutralize these correlations. To test this hypothesis, we ran a nonparametric runs test on the pricing differences over the sample

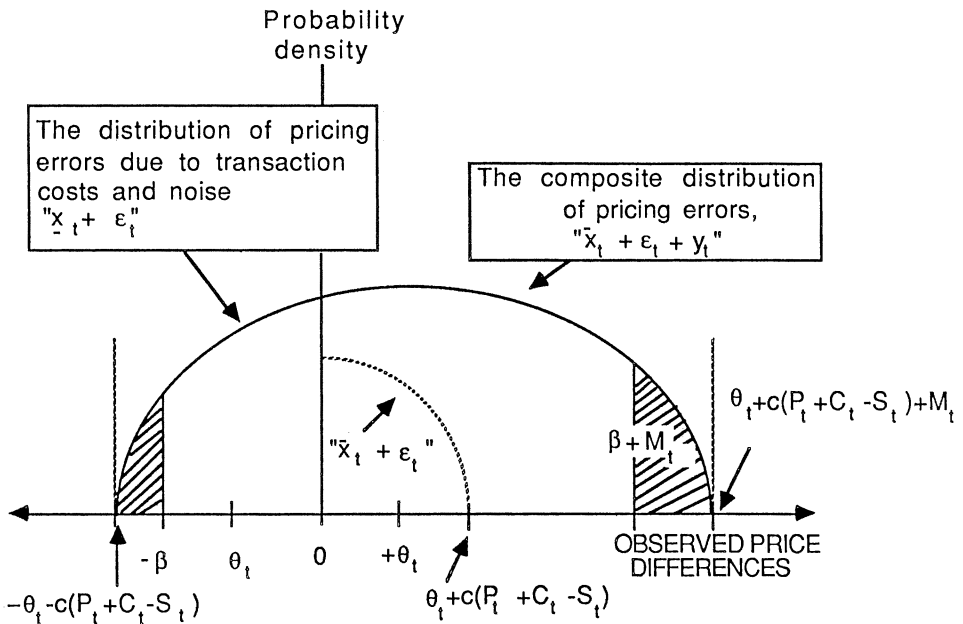


Figure 2. Daily closing price quotations from the *Wall Street Journal* of (Prime + Score – Stock). This figure shows price differences for Exxon over the period September 10, 1985–June 30, 1987. The horizontal axis begins counting with September 10, 1985 as day zero.

What matters for our analysis is whether these differences exhibit any systematic bias. In particular, are primes and scores overpriced relative to the underlying stock? Table I presents evidence that they are overpriced. For each company, the mean of prime and score prices exceeds the stock price. Perhaps more important, the data suggest that this overpricing is not an isolated phenomena; in every case, a large fraction of the observations exhibits higher prices for the prime and score than for the stock. For example, on 319 of 426 days the Exxon prime and score prices exceeded the stock price, with an average positive difference on those days of 73 cents.⁸

Although these data present evidence of overpricing, these price differences need not be significant. Due to transaction costs, price differentials can prevail without violating market efficiency. Noise could also partially explain the price deviations. Using the methodology developed in the previous section, we can test whether these factors explain the observed price discrepancies.

As a first step, we calculated the number of price deviations exceeding the

period. For all companies, the null hypothesis of zero serial correlation could not be rejected at the 95% confidence level. This supports the maintained i.i.d. hypothesis.

⁸ Because we use closing data from the *Wall Street Journal*, our concern is that the measurement error components will not satisfy our boundedness assumption. (See Appendix.) Such a problem, if it existed, would invalidate our testing methodology. To investigate this possibility, a sample of transaction prices for all traded primes and scores for one week in June 1988 was studied. The evidence reveals a pricing pattern almost identical to that reported above. This suggests that end-of-day trading anomalies are not responsible for the price behavior documented in this study. We are grateful to James MacBeth and David Emmanuel for providing the data and analysis on this issue.

Table I
Prime Plus Score Less Stock Price
Differences

This table reports prime plus score less stock price differences obtained from daily price observations as reported in the *Wall Street Journal* over the time period September 10, 1985–June 30, 1987.

	Dollar Differences	Dollar Differences > 0
Exxon		
Number	426	319
Mean	0.473	0.731
Std. Dev.	0.626	0.457
Max. Value	3.000	3.000
Min. Value	−3.000	0.120
American Home Products		
Number	114	84
Mean	0.580	0.911
Std. Dev.	0.765	0.592
Max. Value	2.625	2.625
Min. Value	−1.000	0.125
Bristol-Myers		
Number	92	71
Mean	0.806	1.234
Std. Dev.	1.050	0.713
Max. Value	3.500	3.500
Min. Value	−1.875	0.125
DuPont		
Number	79	57
Mean	0.782	1.303
Std. Dev.	1.200	0.971
Max. Value	3.875	3.875
Min. Value	−1.625	0.125
Merck		
Number	61	39
Mean	0.887	1.859
Std. Dev.	1.688	1.257
Max. Value	5.500	5.500
Min. Value	−2.375	0.125

theoretical 5% right-hand tail cutoff (as explained in the previous section with $M = 0$). Excessive observations above this level may represent arbitrage opportunities. We then tested to see whether the number of violations is significantly larger than 5%. The results in Table II indicate a high degree of overpricing. All five deviations are significant at the 90% confidence level. The prices of Exxon primes and scores, for example, violate this upper pricing bound 25.3% of the time, with AHP primes and scores above the bound 42.1% of the time. Since our tests are designed to be conservative, these large percentage deviations are strong rejections of the null hypothesis.

Table II
Prime Plus Score Less Stock Price Differences Exceeding Cutoff Level

This table reports prime plus score less stock price differences which exceed the 95 percent level determined from the left-hand sample distribution by symmetry. These statistics are based on daily price observations as reported in the *Wall Street Journal* over the time period September 10, 1985–June 30, 1987.

Company	Exxon	American Home Products	Bristol-Myers	DuPont	Merck
Sample Size	426	114	92	79	61
Number of Violations	108	48	18	25	18
Percent of Violations Which Exceed 95% Level	25.3%*	42.1%*	19.6%*	31.6%*	29.5%*
Mean	0.368	0.668	0.640	1.156	1.368
Std. Dev.	0.332	0.476	0.520	0.871	0.858
Max. Value	2.126	1.982	1.925	2.880	3.875
Min. Value	0.006	0.103	0.045	0.013	0.375

* Significant at the 90 percent confidence level.

Because our testing procedure uses the data to estimate implicitly the level of transaction costs and noise, one question that naturally arises is how our transaction cost estimates compare with those used in other studies. One difficulty in comparing transaction cost results is that our adjustment also includes the effects of nonsynchronous trading and typographical errors. Nonetheless, the median of our left-hand distribution provides an upper-bound estimate of the average level of transaction costs.⁹ In every case, these medians are less than 0.5% of the stock's value. By comparison, Sweeney (1988) estimates that actual out-of-pocket trading costs range from 0.1% for floor brokers to 0.4% for private investors. Our average transaction cost results (which include a positive bias due to noise) are thus comparable to those employed in similar studies.

Our approach to estimating transaction costs differs from previous studies in that, to incorporate noise, we use a transaction cost level at the 95% fractile, rather than at the median. These numbers are much higher, ranging from 0.75% to 1.5% of the stock's value. This suggests that factors such as bid/ask spreads and nonsynchronous trading allow observed prices to deviate widely from theoretical bounds. Despite these large price adjustments, however, the data indicate significant overpricing of the prime and score for every firm in our sample. Since the prime and score sell at a premium to the stock, it may be possible to exploit these price differences to make an arbitrage profit.

⁹ To simplify the presentation, we provide the argument for the expected transaction cost to the left of zero, rather than for medians. For a symmetric distribution, these arguments are identical. The average of the deviations to the left of zero is a consistent estimate of $E[\hat{\delta}_t | \hat{\delta}_t \leq 0]$. Consider the following sequence of equalities and inequalities:

$$\begin{aligned}
 E[\hat{\delta}_t | \hat{\delta}_t \leq 0] &= E[\delta_t + \epsilon_t | \hat{\delta}_t \leq 0] \\
 &= E[\bar{x}_t + \bar{x}_t + y_t + \epsilon_t | \hat{\delta}_t \leq 0] \geq E[\bar{x}_t | \hat{\delta}_t \leq 0] + E[\epsilon_t | \hat{\delta}_t \leq 0] \geq E[\bar{x}_t | \hat{\delta}_t \leq 0]
 \end{aligned}$$

since $\bar{x}_t, y_t \geq 0$, and $E[\epsilon_t | \delta_t + \epsilon_t \leq 0] \leq 0$ since the distribution of δ_t is skewed to the right of zero, and ϵ_t is symmetric around zero. This completes our argument.

The previous section demonstrated that arbitraging the price differences between the stock and prime plus score requires additional costs, not considered above, due to either creating primes and scores through the trust or selling short. We first consider the cost of transacting through the trust. The trust charges a deposit fee for each share tendered in this process. (The actual fees were obtained from the appropriate prospectus.) Since this fee is a decreasing function of the trade size, the cost of implementing this strategy differs with the scale of the transaction.

Table III gives the results of our tests with the actual deposit fees incurred for various transaction sizes for each firm in our sample. The table indicates that, without exception, we cannot reject the arbitrage-based boundaries at the 90% confidence level for trades of 1000 shares or less. Given the price taking assumption implicit in our boundary conditions, these transaction sizes provide the best tests for arbitrage profits. Here, the evidence is uniformly consistent with no violation of the arbitrage-based boundaries. Even though price differentials exist, the market appears to be efficient.

For larger transactions, however, arbitrage opportunities may exist if purchasing the stock could be accomplished without significant price effects. For DuPont, trades of at least 1000 shares and, for AHP, trades of at least 5000 shares are required before the arbitrage boundaries are significantly violated. For Exxon, B-M, and Merck, the data dictate that trades of at least 10,000 shares (or block trades) are required. Because block trades tend to have price effects (see Burdett and O'Hara (1987), Easley and O'Hara (1987), and Holthausen, Leftwich, and Mayers (1987)), these large trade sizes may preclude making actual arbitrage profits. We consider this issue in more detail later in the paper.

One difficulty with going through the trust to arbitrage is that it requires the creation of additional primes and scores. However, since each trust is limited to holding no more than 5% of the outstanding stock, there is a limit on the number of primes and scores the trust can create. Once a trust is closed, this method of arbitrage is infeasible. An alternative arbitrage strategy is to sell primes and scores short. To guarantee the profit, this position must be held until the termination of the trust when the prime and score are reconverted to the stock. As with any listed security, the proceeds from a short sale must be held in a margin account (for the life of the short sale). In the case of primes and scores, this constraint is particularly important because the short must be held for the five-year life of the trust.¹⁰

For each firm in our sample we calculated the interest expense of maintaining the short sale over the life of the trust. Because of different expiration dates and times since listing, these costs differ across trusts. To determine the relevant interest rate, we used the *Wall Street Journal* to obtain daily prices of the Treasury security maturing closest to each trust's expiration date. Considering coupons, an internal rate of return on the bonds was calculated based on the average of the bid and ask price quotes. The short sale cost was then calculated based on this yield. Not surprisingly, these costs are quite large. For each firm in

¹⁰ A related concern is the difficulty introduced by the price tick rule prohibiting short sales whenever the price exhibits a down-tick. Since this may seriously impede the ability to exploit any temporary price discrepancies, this method of arbitrage may be severely limited.

Table III
Violations of Arbitrage Pricing Bounds

This table reports percent violations of arbitrage bounds which exceed the 95 percent level plus trust exchange fees. The 95 percent level is determined from the left-hand sample distribution assuming symmetry. The transactions are ordered by size of deposit fee. The calculations are based on daily price observations from the *Wall Street Journal* over the time period September 10, 1985–June 30, 1987.

Transaction Size in Shares	100– 1000	1001– 5000	5001– 10,000	10,001– 50,000	50,001– 100,000	100,001– 250,000	250,001– 1,000,000	>million
	American Home Products							
Sample Size	114	114	114	114	114	114	114	114
Percent Violations (Dollar Data)	7.02	10.53	14.04*	18.42*	21.05*	21.05*	27.19*	30.70*
Percent Violations (Normalized Data ^a)	7.02	10.53	14.04*	18.42*	19.30*	21.05*	27.19*	30.70*
Bristol-Myers								
Sample Size	92	92	92	92	92	92	92	92
Percent Violations (Dollar Data)	10.87	10.87	10.87	10.87	10.87	14.13	14.13	14.13
Percent Violations (Normalized Data ^a)	10.87	10.87	10.87	11.96	14.13	14.13	16.30*	16.30*
DuPont								
Sample Size	79	79	79	79	79	79	79	79
Percent Violations (Dollar Data)	12.66	16.46*	17.72*	17.72*	18.99*	21.52*	24.05*	26.58*
Percent Violations (Normalized Data ^a)	13.92	16.46*	16.46*	16.46*	20.25*	21.52*	21.52*	24.05*
Merck								
Sample Size	61	61	61	61	61	61	61	61
Percent Violations (Dollar Data)	6.56	8.20	9.84	9.84	11.48	16.39	21.31*	24.59*
Percent Violations (Normalized Data ^a)	4.92	8.20	9.84	9.84	11.48	14.75	19.67*	22.95*
Exxon ^b								
Sample Size	426	426	426	426	426	426	426	426
Percent Violations (Dollar Data)	1.64	3.52	5.87	7.98	10.56*	11.03*	11.03*	11.03*
Percent Violations (Normalized Data ^a)	3.99	7.28	9.86*	12.91*	15.49*	17.14*	17.14*	19.25*

* Significant at the 90% confidence level. The cutoff level is 9.08% for Exxon, 14.04% for AHP, 15.38% for Bristol-Myers, 16.45% for DuPont, and 18.57% for Merck.

^a Normalized Data is obtained by dividing the absolute price differences by the stock price.

^b Deposit fees for Exxon are the same for trades between 100 and 5000 shares.

our sample, the average interest lost in short sales ranged from 28 to 33% of the prime plus score value.

We incorporated these short sale costs into our arbitrage bounds and tested for violations. The results are conclusive: there are no violations considering short selling costs to the prime and score. Because the cost of maintaining the short position is so high, no observed price ever fell outside our upper bound. In the case of Exxon, for example, the cost of short selling ranges from 42 to 21% of the stock's value over our sample period. The enormity of these costs means that prime and score prices can deviate widely from the stock's price without ever violating the arbitrage boundary.

This illustrates an important problem with short sale constraints. In most short sales, the position is reversed in a matter of days, so any constraint on the proceeds is not unduly costly. With primes and scores, however, a long-term position must be taken to lock in the price disparities. For traders subject to the proceeds constraint, this is far too expensive to maintain and thus effectively precludes the use of short sales for any long-run arbitrage use. Since the alternative arbitrage strategy is limited by constraints on the size of the trust, this suggests that once a trust is closed the market may be ineffective in bringing prime and score prices in line with stock prices.

Thus far we have demonstrated that, for traders subject to standard trust fees or to the proceeds constraint on short sales, the opportunities for arbitrage profits are limited. For traders not subject to these costs, however, the potential for profits is obviously greater. In particular, Table II indicated that traders could earn arbitrage profits in every firm in our sample if they faced only standard transaction costs. The size of these profits, however, depends crucially on the price effects of trades. Specifically, implicit in this arbitrage profit calculation is the assumption that trades could have been completed at the closing prices used in our sample. If, in the process of trading, the trade itself affects prices, however, then these potential profits may be illusory.

To investigate this issue, we calculated the median daily trading volume in the prime, score, and stock for each firm in our sample. Since price effects are more likely in thin markets, the size of trading volume is a good indicator of the seriousness of this problem. As Table IV indicates, the trading volume of primes and scores is very small relative to that of the stock. Moreover, significant volume differences also exist between the prime and score, with median daily prime volumes ranging from only 6,000 to 22,000 shares.

These figures suggest that attempts to buy or sell large amounts of primes and scores are likely to have significant price effects. Consequently, even for traders not subject to the proceeds constraint on short sales, the price effects of trades may compromise any ability to earn arbitrage profits. In particular, Table III indicated that single trades of at least 10,000 shares were needed to profit in Exxon and Merck. Since the entire daily prime volume in these companies was only 12,000 and 7,000 shares, respectively, such trades at the reported prices are most likely infeasible.

Whether arbitrage opportunities exist, therefore, depends on many factors including the method and scale of operation. What the results in this section

Table IV
Median Daily Trading Volume

This table reports median daily trading volume in thousands of shares as reported in the *Wall Street Journal* over the time period September 10, 1985–June 30, 1987.

Company	All Trading Days				Around Ex-Dividend Dates (± 2 Days)			
	Sample Size	Prime	Score	Stock	Sample Size	Prime	Score	Stock
American Home Products	131	18	89	2943	9	6	46	2725
Bristol-Myers	94	22	60	4070	5	17	38	2966
DuPont	107	6	79	5821	4	29	93.5	4201
Exxon	455	12	82	8827	33	12	75	10056
Merck	67	7	15	4978	5	0	12	5127

demonstrate is that the overpricing of primes and scores is not an illusion: the market values primes and scores more highly than it does the underlying stock. For at least some market participants, therefore, the ability to trade a dividend-based component and an option-based component separately is valuable. What is unclear is why such value should arise. In the next section we analyze the possible origins of this overpricing.

IV. The Valuation of Primes and Scores

Why would investors pay a premium to purchase these securities? Theories of financial innovation suggest several possible answers. (See Miller (1986), Van Horne (1985), and Silber (1975).) In general, these explanations rely on market completeness, transactions costs, or taxes to explain how value can arise through financial innovation. In this section we investigate which, if any, of these three explanations is consistent with the behavior of primes and scores. Our focus is on identifying those factors that lead to the overpricing pattern identified in the previous section.

One hypothesis is that value arises from the completing of markets. If primes and scores provide a way to span states of nature previously unavailable to the investor, then these securities will be more valuable than a standard pricing framework would suggest. In the case of primes and scores, this completeness may arise because the score represents a new security. Although traded options exist for each firm in our sample, these options are all short term; the longest option currently traded on the Chicago Board of Options Exchange (CBOE) has a maturity of nine months. Because the score is a long-term (five-year) European call option, it may allow investors to undertake trading strategies currently unavailable. Splitting a stock into a prime and a score, therefore, potentially creates a new pattern of cash flows not previously available.

A second hypothesis is that the price behavior is due to transaction costs. Specifically, the prime and score may provide a cheaper way to achieve some objective than was previously available. One reason this may occur is that the score provides a means of saving on the costs of dynamic hedging. Specifically,

although it is true that traded long-term options do not exist, it may be possible to replicate a long-term option through a process of dynamic hedging. Consequently, by taking positions in a risk-free bond and the stock, an investor can potentially create a synthetic long-term option, even though such a security does not currently exist on any exchange. One difficulty in doing this, however, is that it involves continual rebalancing of the underlying position. Since this, in turn, involves enormous transaction costs, the theoretical ability to create such an option may be of little value to investors. Even if the long-term characteristic of the score is not unique, the score may still be valuable if it economizes on the transaction costs of dynamic hedging.

A final hypothesis is that the price differential is due to tax effects. These tax effects could arise from either the corporate dividend exclusion rule or because of a tax timing option. The dividend exclusion explanation arises because corporations can exclude 80% of dividend income from income taxes. Because the prime provides a cheaper means of acquiring the dividend stream than does purchasing the stock, the prime may be more valuable to corporate investors. This suggests that the increased value may be due to an increased demand for the prime. The timing option explanation arises because of the differential tax treatment of long-term capital gains. (See Constantinides (1984).) Specifically, suppose that an investor has a short-term gain. During our sample period, if he or she sells the stock, he or she must recognize the entire gain for tax purposes. By selling the prime and keeping the score, however, the investor can postpone realizing any gains until they are treated as capital gains. This timing option explanation implies that overpricing is due to the score and not the prime.

To test these hypotheses, we first consider the behavior of the score. Our volume data (see Table IV) indicate much greater trading in the score than in the prime, providing preliminary evidence consistent with the score being the more desirable component. To provide additional evidence, however, we need to examine actual price behavior.

Under either the market completeness or transaction cost hypothesis, the score price should reflect increased value. One way to test these theories is to compare actual score prices with those option values predicted by a model based on complete markets and without transaction costs. The Black-Scholes option model satisfies these conditions. Under the assumptions of frictionless markets (i.e., no taxes, margin requirements, transaction costs), constant interest rates (r per unit time), lognormally distributed prices with volatility parameter σ , and known dividends of d_1 , d_2 dollars at times t_1 , t_2 , respectively, a European call with an exercise price K and maturity T has value

$$C_t = [S_t - d_1 e^{-r(t_1-t)} - d_2 e^{-r(t_2-t)}]N(h) - Ke^{-r(T-t)}N(h - \sigma\sqrt{T-t}), \quad (4)$$

where $h = \log[(S_t - d_1 e^{-r(t_1-t)} - d_2 e^{-r(t_2-t)})/Ke^{-r(T-t)}]/\sigma\sqrt{T-t} + 1/2\sigma\sqrt{T-t}$.

The frictionless market hypothesis can be partially relaxed in that Scholes (1976) and Heath and Jarrow (1987) have shown that taxes and margin requirements, respectively, do not change the formula. Transaction costs, however, affect the option's valuation through the dynamic hedge. Since hedging costs are not incorporated into a historical volatility estimation procedure, using a histor-

ical volatility measure provides one method to estimate a zero transaction cost value for the score. Alternatively, Leland (1985) and Jarrow and Wiggins (1988) argue that hedging costs are partially captured through an implicit volatility estimation procedure because, under proportional transaction costs, the standard Black–Scholes hedge with an adjusted volatility still approximates the option's value. The volatility is adjusted to reflect both the round-trip transaction costs (measured as a fraction of the dollar volume of transactions) and the discrete hedging interval; see Leland (1985, p. 1289). Thus, the difference between estimated score prices with historic volatilities and those estimated with implicit volatilities provides a rough estimate of the transaction costs involved in the dynamic hedge.

Valuing the score utilizing both historic and implicit volatilities will provide some evidence concerning the market completeness and transaction cost hypotheses. If the historic volatility-based model accurately predicts score values, then neither the market completeness nor the transaction cost hypothesis will be consistent with the evidence. Conversely, if the implicit volatility model predicts the score's price better than the historic, then the overpricing could be due to the transaction cost hypothesis. It could, however, also be due to a violation of the lognormality assumption since implicit volatilities would partially adjust for misspecifications of the stock price process as well.

To estimate option values, we obtained historic dividends on the five stocks as reported in Moody's Dividend Record. These dividends were projected forward five years using a simple extrapolation of past dividend performance and ex-dividend dates. Historic volatilities were calculated using a rolling, sixty-day sample standard deviation. Implicit volatilities were calculated numerically, based on the score price of the preceding day.

Table V presents the results of our tests. The data indicate that on average the implicit volatility model accurately prices the score but the historic volatility model does not. In all five cases, the implicit volatility option prices are remarkably close to actual score values. These differences are of magnitudes comparable to those observed in previous studies based on implicit volatilities. (See Whaley (1982).) Based on this evidence, it appears that one cannot reject either the transaction cost hypothesis nor the market completeness hypothesis since historic volatilities do not accurately predict score values. In fact, using historic volatilities, score pricing errors as large as 25% are observed. To statistically test these hypotheses, we ran the following regression for each firm in our sample:

$$C_t = a_0 + a_1 \hat{C}_t + \mu_t, \quad (5)$$

where C_t represents the market price and $\hat{C}_t$ represents the model price.

These tests are reported in Table VI. The null hypothesis of ($a_0 = 0$) and ($a_1 = 1$) is rejected in all cases for the historic volatility. This hypothesis is rejected using implicit volatilities only for Merck, AHP, and B-M.¹¹ The failure of the

¹¹ Additional diagnostic analysis on the pricing errors was performed by running the regression

$$(C_t - \hat{C}_t)/\hat{C}_t = a_0 + a_1 \hat{\sigma}_t + \mu_t$$

for each score in our sample. These regressions confirm that the pricing error is statistically correlated to the historic volatility estimate but not to the implicit.

Table V**Score Prices and Black-Scholes Option Values**

Values were calculated with different volatility measures as reported in the *Wall Street Journal* over the time period September 10, 1985–June 30, 1987.

	Actual Score Price	B-S Price Historic Vol. ^a	B-S Price Imp. Vol. ^b	B-S Price CBOE Vol. ^c
Exxon				
Number	454	446	444	408
Mean	14.44	13.87	14.49	12.32
Std. Dev.	8.81	7.27	8.82	7.91
Max. Value	34.50	31.83	34.86	30.21
Min. Value	5.50	5.11	4.64	3.19
American Home Products				
Number	132	130	129	122
Mean	18.16	21.07	18.22	16.24
Std. Dev.	2.18	2.09	2.19	2.82
Max. Value	22.88	25.43	22.65	21.91
Min. Value	10.63	16.02	10.32	9.11
Bristol-Myers				
Number	94	93	92	89
Mean	30.33	23.72	30.33	25.95
Std. Dev.	2.35	2.12	2.41	3.12
Max. Value	35.75	28.93	36.06	31.17
Min. Value	26.00	18.14	25.29	17.54
DuPont				
Number	105	106	103	106
Mean	36.44	31.50	36.43	34.83
Std. Dev.	6.63	5.07	6.51	6.62
Max. Value	45.50	41.21	45.73	44.21
Min. Value	24.25	20.40	24.35	17.11
Merck				
Number	66	66	64	64
Mean	46.17	49.08	46.19	44.82
Std. Dev.	2.75	4.83	3.10	4.64
Max. Value	50.63	58.46	52.31	51.80
Min. Value	41.38	40.65	40.50	34.86

^a B-S Price Historic Vol. is based on a rolling sixty-day sample standard deviation of price returns.

^b B-S Price Imp. Vol. is based on an implicit volatility from the preceding day.

^c B-S Price CBOE Vol. is based on the implicit volatility from the nearest at-the-money Chicago Board Options Exchange call.

historic volatility model is consistent with transaction costs playing an important role in the pricing of the score. If these prices consistently underestimated the score value, the evidence would be stronger. Unfortunately, in two cases (AHP and Merck), the model overpredicts the score value, a result inconsistent with this argument but consistent with a violation of the lognormality hypothesis.

Table VI
Tests of Score Pricing Models

Tests based on three different measures of volatilities using closing prices as reported in the *Wall Street Journal* over the time period September 10, 1985–June 30, 1987.

$C_t = a_0 + a_1 \hat{C}_t$						
Sample Size	Volatility Estimation	a_0	$t(a_0)$	a_1	$t(a_1)$	R^2
Exxon						
$N = 445$	Hist. Vol.	-2.1600	-14.709	1.19793	21.104	0.974
$N = 407$	CBOE Imp. Vol.	1.1710	5.939	1.09177	6.812	0.942
$N = 443$	Score Imp. Vol.	0.0642	1.400	0.99697	-1.121	0.997
American Home Products						
$N = 129$	Hist. Vol.	-1.8041	-2.085	0.94796	-1.274	0.808
$N = 120$	CBOE Imp. Vol.	10.6340	12.007	0.46539	-9.964	0.387
$N = 128$	Score Imp. Vol.	1.1793	3.306	0.93586	-3.300	0.948
Bristol-Myers						
$N = 92$	Hist. Vol.	19.2976	7.627	0.46492	-5.035	0.174
$N = 88$	CBOE Imp. Vol.	25.7244	12.779	0.17218	-10.748	0.054
$N = 91$	Score Imp. Vol.	2.0502	2.086	0.93134	-2.125	0.902
DuPont						
$N = 105$	Hist. Vol.	4.0289	1.213	1.00419	0.040	0.474
$N = 104$	CBOE Imp. Vol.	5.0878	1.876	0.88347	-1.529	0.566
$N = 102$	Score Imp. Vol.	0.1858	0.289	0.99582	-0.241	0.970
Merck						
$N = 65$	Hist. Vol.	12.7317	1.796	0.66657	-2.320	0.252
$N = 63$	CBOE Imp. Vol.	19.6752	2.690	0.57667	-2.608	0.169
$N = 63$	Score Imp. Vol.	10.3811	3.861	0.77418	-3.888	0.741

C_t is the market price of the score at time t .

$\hat{C}_t$ is the model price for the score at time t .

$t(a_0)$ represents the t -statistic for the null hypothesis of $a_0 = 0$.

$t(a_1)$ represents the t -statistic for the null hypothesis of $a_1 = 1$.

Hist. Vol. is a sixty-day rolling sample standard deviation of price returns.

CBOE Imp. Vol. is an implicit volatility from the nearest at-the-money CBOE call.

Score Imp. Vol. is an implicit volatility based on the score price of the preceding day.

To further differentiate the transaction cost hypothesis from the market completeness hypothesis, we examined alternative data. In particular, since all firms in our sample have CBOE traded options, we use these prices to obtain a third estimate of the implicit volatility. If the implicit volatility reflects the cost of a dynamic hedge over the option's life, implied volatilities from CBOE options should imply lower option values than the implied volatilities based on the score

price. Further, they should exceed those obtained from the historic volatility. To calculate these values, the nearest at-the-money call option's closing price was obtained from the *Wall Street Journal*. We then numerically calculated its implicit volatility for the same day the score price was estimated.

The results from Table V provide additional but weak evidence consistent with the transaction cost hypothesis. In all five cases, the Black-Scholes prices with the CBOE volatilities are less than the Black-Scholes prices with the implicit volatility estimated from the score itself. Further, in all but three cases (Exxon, AHP, and Merck) the CBOE volatility Black-Scholes prices exceed the historic volatility prices. Two of these cases (Merck and AHP), however, correspond to those cases where the BS price based on historic volatilities already exceeded the score-based implied volatility BS prices. These discrepancies could be due to either the standard measurement error problems inherent in using closing prices as discussed in the preceding section or a violation of the lognormality assumption.

Our analysis thus far has examined whether a market completeness or a transaction cost hypothesis is consistent with the prime and score price behavior. A third possibility is that the increased value is due to tax effects. As noted earlier, these tax effects could arise either from the corporate dividend exclusion rule or from a tax-timing option. One way to test these theories is to see whether the prices of primes and scores were influenced by the recent tax law change. The tax change both lowered the amount of dividends qualifying for exclusion (from 85 to 80%) and raised the tax rate on capital gains. If the dividend exclusion argument holds, this first change should have reduced the attractiveness of the prime. Similarly, if the tax-timing explanation is correct, the second change should have reduced the attractiveness of the score. A decrease in the differential of prime and score prices around the tax change, therefore, may give evidence of some tax-induced behavior.

One difficulty in testing this hypothesis is that it requires time-series data spanning periods both before and after the tax change. Since most securities in our sample are less than one year old, only Exxon provides sufficient data to test this hypothesis. We regressed the observed daily price discrepancies on a constant plus a dummy variable D_t , where D_t is one after January 1, 1987 and zero otherwise. The results obtained are

$$\begin{array}{rcl} [P_t + C_t - S_t] & = & 0.02074 + 0.001566D_t. \\ \text{(standard errors)} & & (0.02805) \quad (0.5330) \end{array} \quad (6)$$

The R^2 was 0.00 for a sample size of 443. There appears to be no effect on prices around the date the tax law changed. Based on this limited data, therefore, it appears that tax effects are not responsible for the overpricing. Yet this is a weak test since the unanticipated change in valuation due to the tax law change, at year's end, could have been very small, even if the prime or score was overvalued before an *announcement* of the tax law change.

As a further test of the dividend exclusion hypothesis, we segmented our data around ex-dividend days (± 2 days) to test for any dividend-induced volume differences. In particular, if primes are purchased only to capture the dividend, then the volume of prime trading around ex-dates should reflect this increased

activity. Table IV indicates (if anything) the opposite effect. The average trading volume around ex-dividend dates is less than the normal value given in Table IV. The dividend exclusion hypothesis is not supported by these observations.

Our results, although encouraging, are inconclusive. Based on weak evidence, the transaction cost hypothesis appears to be the leading explanation for prime and score price behavior. However, we also cannot reject either the market completeness hypothesis or the tax hypothesis. Additional research is needed along these lines.

V. Conclusions

The introduction of primes and scores gives graphic support to the argument (see Van Horne (1985) and Miller (1986)) that the presence of market imperfections motivates the process of financial innovation. By addressing inefficiencies in the market due to factors such as taxes, transaction costs, and incompleteness, financial innovations provide value both to their developers and to the financial markets as a whole. As we have argued in this paper, dividing equities into prime and score components results in their value together exceeding that of the underlying stock. This price disparity is stable because regulatory constraints on short sales and trust size may make transaction costs so large as to preclude arbitrage. Hence, just as market imperfections brought these instruments into existence, so, too, do they allow primes and scores to flourish.

Perhaps it is fitting that yet another market imperfection (taxes or, more specifically, the IRS) is responsible for the demise of primes and scores. Although the new tax ruling effectively ends the introduction of additional primes and scores, the success of these instruments suggests that similar innovations should be expected to appear. As Van Horne (1985) noted, a successful innovation often spawns other innovations as promoters enter the market in response to profit opportunities. The recently proposed unbundled stock units suggest that just such a second phase of development may be occurring. These new securities involve segmenting equities into three or more parts designed to isolate the risk return characteristics of the underlying stock. Although it is too soon to tell which specific new securities will prove successful, our analysis here demonstrates that it may, indeed, be possible to create value simply by splitting a security into different parts.

Appendix: Statistical Methodology for Testing Arbitrage Bounds

The given are a probability space $(\Omega, F, \text{prob}(\cdot))$, where Ω is a state space, F a σ -algebra, and $\text{prob}: F \rightarrow [0, 1]$ a complete probability measure, and two random variables, $\delta_t: \Omega \rightarrow R$ and $\epsilon_t: \Omega \rightarrow R$.

Define

$$\hat{\delta}_t \equiv \delta_t + \epsilon_t,$$

$$\underline{x}_t \equiv \delta_t 1_{|\delta_t < 0|},$$

$$(\bar{x}_t + y_t) \equiv \delta_t 1_{|\delta_t \geq 0|},$$

where

$$1_{\{\delta_t < 0\}} = \begin{cases} 1 & \text{if } \delta_t < 0, \\ 0 & \text{otherwise,} \end{cases}$$

$$1_{\{\delta_t \geq 0\}} = 1 - 1_{\{\delta_t < 0\}}.$$

Assume the following six conditions:

- (A1) $-A_t \leq \delta_t \leq A_t + M_t$, where $A_t, M_t \geq 0$ are constants,
- (A2) $\text{prob}(\underline{x}_t \leq -\beta) = \text{prob}(\bar{x}_t \geq \beta)$ for all $\beta \geq 0$,
- (A3) $0 \leq y_t \leq M_t$ for $M_t \geq 0$ a constant,
- (A4) $\text{prob}(\epsilon_t \leq -\beta) = \text{prob}(\epsilon_t \geq \beta)$ for all $\beta \geq 0$,
- (A5) $-\theta_t \leq \epsilon_t \leq \theta_t$ for $\theta_t \geq 0$ a positive constant,
- (A6) ϵ_t are statistically independent of $\{\underline{x}_t, \bar{x}_t, y_t\}$.

Remarks:

1. $-A_t \leq \underline{x}_t \leq 0$ with probability one, by (A1).
2. $0 \leq \bar{x}_t \leq A_t$. This follows from (A2) and (1).
3. $\text{prob}(\underline{x}_t + \epsilon_t \leq -\beta) = \text{prob}(\bar{x}_t + \epsilon_t \geq \beta)$ for all $\beta \geq 0$.

Proof:

$$\begin{aligned} & \text{prob}(\underline{x}_t + \epsilon_t \leq -\beta) \\ &= \int_R \text{prob}(\underline{x}_t \leq -\beta - \alpha) d \text{prob}(\epsilon_t \leq \alpha) \quad \text{by (A6)} \\ &= \int_R \text{prob}(\bar{x}_t \geq \beta + \alpha) d \text{prob}(\epsilon_t \leq \alpha) \quad \text{by (A2)} \\ &= \text{prob}(\bar{x}_t - \epsilon_t \geq \beta) \quad \text{by (A6)} \\ &= \int_{[0, A_t]} \text{prob}(-\epsilon_t \geq \beta - \alpha) d \text{prob}(\bar{x}_t \leq \alpha) \quad \text{by (A6)} \\ &= \int_{[0, A_t]} \text{prob}(-\epsilon_t \leq -\beta + \alpha) d \text{prob}(\bar{x}_t \leq \alpha) \quad \text{by (A4)} \\ &= \text{prob}(\bar{x}_t + \epsilon_t \geq \beta). \quad \text{Q.E.D.} \end{aligned}$$

Proof of Lemma 1:

$$\begin{aligned} & \text{prob}(\hat{\delta}_t - M_t \geq \beta) \\ &= \text{prob}(\delta_t + \epsilon_t \geq \beta + M_t) \\ &= \text{prob}(\underline{x}_t + \epsilon_t \geq \beta + M_t \text{ and } \delta_t < 0) \\ & \quad + \text{prob}(\bar{x}_t + y_t + \epsilon_t \geq \beta + M_t \text{ and } \delta_t \geq 0) \end{aligned}$$

$$\begin{aligned}
&\leq \text{prob}(\underline{x}_t + \epsilon_t \geq \beta + M_t) + \text{prob}(\bar{x}_t + y_t + \epsilon_t \geq \beta + M_t) \\
&\hspace{15em} \text{since } y_t \leq M_t, \\
&\leq \text{prob}(\underline{x}_t + \epsilon_t \geq \beta + M_t) + \text{prob}(\bar{x}_t + \epsilon_t \geq \beta). \\
&= \text{prob}(\bar{x}_t + \epsilon_t \geq \beta) \\
&\hspace{15em} \text{since } M_t \geq 0, \quad \underline{x}_t \leq 0, \quad \text{and } \beta > \theta_t \\
&\hspace{10em} \text{imply that } \text{prob}(\underline{x}_t + \epsilon_t \geq \beta + M_t) = 0 \\
&= \text{prob}(\underline{x}_t + \epsilon_t \leq -\beta) \quad \text{by (3)} \\
&= \text{prob}(\delta_t + \epsilon_t \leq -\beta \text{ and } \delta_t < 0) \\
&\hspace{15em} \text{since } -\theta_t \leq \epsilon_t \leq \theta_t \quad \text{and } -\beta \leq -\theta_t. \\
&\leq \text{prob}(\delta_t + \epsilon_t \leq -\beta) = \text{prob}(\hat{\delta}_t \leq -\beta). \quad \text{Q.E.D.}
\end{aligned}$$

Assume that $\{\hat{\delta}_t: t = 1, 2, \dots, n\}$ are independent and identically distributed random variables. This implies, due to (A1) and (A5), that the t subscripts on A_t , M_t , and θ_t can be omitted. Let $\{\hat{\gamma}_t: t = 1, \dots, n\}$ be the order statistics obtained by ranking $\{\hat{\delta}_t: t = 1, \dots, n\}$ from smallest to largest.

Let

$$\begin{aligned}
F(\alpha) &\equiv \text{prob}(\hat{\delta}_t \leq \alpha), \\
F_n(\alpha) &\equiv |\{i \in \{1, \dots, n\}: \hat{\delta}_i \leq \alpha\}|/n, \quad \text{and} \\
\beta_q &\equiv \inf\{\alpha: F(\alpha) \geq q\} \quad \text{for all } q \in [0, 1].
\end{aligned}$$

From Mood, Graybill, and Boes (1974, p. 513),

$$\text{prob}[\hat{\gamma}_j \leq \beta_q \leq \hat{\gamma}_k] = \sum_{i=j}^{k-1} \binom{n}{i} q^i (1-q)^{n-i} \quad \text{for any } q \in [0, 1].$$

To obtain a confidence interval, we transform the null hypothesis to its equivalent form in terms of order statistics. For $\beta_q \leq -\theta$,

$$1 - F(-\beta_q + M) - F(\beta_q) \leq 0 \quad \text{if and only if } \beta_{(1-q)} \leq -\beta_q + M.$$

Define $\hat{\gamma}_l, \hat{\gamma}_h, \hat{\gamma}_L, \hat{\gamma}_H$ such that, for a fixed $\alpha \in [0, 1]$,

$$\text{prob}[\hat{\gamma}_l \leq \beta_q \leq \hat{\gamma}_h] = \sum_{i=l}^{h-1} \binom{n}{i} q^i (1-q)^{n-i} = 1 - \alpha$$

and

$$\text{prob}[\hat{\gamma}_L \leq \beta_{1-q} \leq \hat{\gamma}_H] = \sum_{i=L}^{H-1} \binom{n}{i} (1-q)^i (q)^{n-i} = 1 - \alpha.$$

Next, form the test statistic $Q(n) \equiv -\hat{\gamma}_l + M - \hat{\gamma}_L$. If $Q(n) > 0$, accept the null hypothesis; otherwise, reject it.

The idea underlying this test statistic is explained as follows. If $-\hat{\gamma}_l + M < \hat{\gamma}_L$, then the $(1 - \alpha)$ percent confidence interval for $-\beta_q + M$ is strictly below

the $(1 - \alpha)$ percent confidence interval for β_{1-q} . As shown next, this rejects the null hypothesis with a confidence level of at least $(1 - \alpha)$ percent. Indeed,

$$\begin{aligned} \text{prob}[Q(n) < 0] &= \text{prob}[-\hat{\gamma}_l + M < \hat{\gamma}_L] \\ &= \text{prob}[-\hat{\gamma}_l + M < \hat{\gamma}_L \text{ and } \beta_q < \hat{\gamma}_l] \\ &\quad + \text{prob}[-\hat{\gamma}_l + M < \hat{\gamma}_L \text{ and } \beta_q \geq \hat{\gamma}_l] \\ &\leq \text{prob}[\beta_q < \hat{\gamma}_l] + \text{prob}[-\hat{\gamma}_l + M < \hat{\gamma}_L \text{ and } -\beta_q + M \leq -\hat{\gamma}_l + M]. \end{aligned}$$

Since, under the null hypothesis ($\beta_{(1-q)} \leq -\beta_q + M$), we have

$$\leq \text{prob}[\beta_q < \hat{\gamma}_l] + \text{prob}[\beta_{(1-q)} < \hat{\gamma}_L] = \alpha/2 + \alpha/2 = \alpha,$$

hence, the confidence level for this test is greater than $(1 - \alpha)$ percent.

Proof of Lemma 2: First, note that, for large n ,

$$\begin{aligned} \text{prob}[\hat{\gamma}_l \leq \beta_q \leq \hat{\gamma}_h] &= 1 - \alpha \approx \Phi(nq + Z_{\alpha/2} \sqrt{nq(1-q)}) - \Phi(nq - Z_{\alpha/2} \sqrt{nq(1-q)}), \\ \text{prob}[\hat{\gamma}_L \leq \beta_{1-q} \leq \hat{\gamma}_H] &= 1 - \alpha \approx \Phi(n(1-q) + Z_{\alpha/2} \sqrt{nq(1-q)}) \\ &\quad - \Phi(n(1-q) - Z_{\alpha/2} \sqrt{nq(1-q)}). \end{aligned}$$

Hence,

$$l = nq - Z_{\alpha/2} \sqrt{nq(1-q)} \quad \text{and} \quad L = n(1-q) - Z_{\alpha/2} \sqrt{nq(1-q)}.$$

Find q such that $F_n(\hat{\gamma}_l) = 0.05$; i.e., $l/n = 0.05$ or $l = (0.05)n$. Substitution (and algebra) yields $[n + Z_{\alpha/2}^2]q^2 - (Z_{\alpha/2}^2 + 0.1n)q + n(0.05)^2 = 0$. The solution is $q = ([Z_{\alpha/2}^2 + 0.1n] + Z_{\alpha/2}^2[Z_{\alpha/2}^2 + 0.19n]^{1/2})/2[n + Z_{\alpha/2}^2]$. Under this q , $F_n(\hat{\gamma}_l) = 0.05$, so $-\hat{\beta} = \hat{\gamma}_l$. Hence,

$$\begin{aligned} \alpha &\geq \text{prob}[Q(n) < 0] = \text{prob}[-\hat{\gamma}_l + M < \hat{\gamma}_L] = \text{prob}[\hat{\beta} + M < \hat{\gamma}_L] \\ &= \text{prob}[F_n(\hat{\beta} + M) < F_n(\hat{\gamma}_L)] = \text{prob}[F_n(\hat{\beta} + M) < L/n] \\ &= \text{prob}[1 - F_n(\hat{\beta} + M) > (n - L)/n]. \end{aligned}$$

REFERENCES

- Ball, C., 1988, Estimation bias induced by discrete security prices, *Journal of Finance* 43, 841-865.
 Bhattacharya, M., 1983, Transaction data tests of efficiency of the Chicago Board Options Exchange, *Journal of Financial Economics* 12, 161-185.
 Bookstaber, R., 1981, Observed Option mispricing and the nonsimultaneity and option quotations, *Journal of Business* 54, 141-155.
 Burdett, K. and M. O'Hara, 1987, Building blocks: An introduction to block trading, *Journal of Banking and Finance* 11, 193-212.
 Constantinides, G. M., 1984, Optimal stock trading with personal taxes: Implications for prices and the abnormal January returns, *Journal of Financial Economics* 13, 65-90.

- Cox, J. and M. Rubinstein, 1985, *Options Markets* (Prentice Hall, Englewood Cliffs, NJ).
- Easley, D. and M. O'Hara, 1987, Price, trade size, and information in securities markets, *Journal of Financial Economics* 19, 69–90.
- Foster, F. D. and G. S. Oldfield, 1987, Ex-dividend day stock price adjustment: Taxes and transaction costs, Unpublished manuscript, Cornell University.
- Heath, D and R. Jarrow, 1987, Arbitrage, continuous trading, and margin requirements, *Journal of Finance* 42, 1129–1142.
- Holthausen, R., R. Leftwich, and D. Mayers, 1987, The effect of large block transactions on security prices: A cross-sectional analysis, *Journal of Financial Economics* 19, 237–268.
- Jarrow, R. and J. Wiggins, 1988, Option pricing and implicit volatilities: A review and a new perspective, *Journal of Economic Surveys*, Forthcoming.
- Leland, H., 1985, Option pricing and replication with transaction costs, *Journal of Finance* 40, 1285–1302.
- Miller, M., 1986, Financial innovation: The last twenty years and the next, *Journal of Financial and Quantitative Analysis* 21.
- Mood, A., F. Graybill, and D. Boes, 1974, *Introduction to the Theory of Statistics*, 3rd ed. (McGraw-Hill, New York).
- Phillips, S. M. and C. Smith, 1980, Trading costs for listed options: The implications for market efficiency, *Journal of Financial Economics* 8, 179–201.
- Scholes, M., 1976, Taxes and the pricing of options, *Journal of Finance* 31, 319–332.
- Silber, W., 1975, *Financial Innovation* (Lexington Books, Lexington, MA).
- Sweeney, R. J., 1988, Some new filter rule tests: Methods and results, *Journal of Financial and Quantitative Analysis* 23, 285–300.
- Van Horne, J. C., 1985, Of financial innovations and excesses, *Journal of Finance* 40, 621–631.
- Whaley, R. E., 1982, Valuation of American call options on dividend-paying stocks, *Journal of Financial Economics* 10, 29–58.