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The Price Elasticity of Demand for Common Stock

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ABSTRACT

We study the price elasticity of demand for the common stock of an individual corporation. Despite the prevelance of assumptions that demand is perfectly elastic, there is little if any direct evidence in the literature to either support or reject that contention. Consistent with the notion of finite price elasticities, we find that the announcement of primary stock offerings by regulated firms depresses their stock prices and little if any evidence that this decline is the result of adverse information about future cash flows. Attempts to relate offer announcement effects directly to possible determinants of price elasticities, however, are inconclusive.

THE PURPOSE OF THIS study is to examine whether the price elasticity of demand for the common stock of any individual corporation is measurably finite. A common assumption in finance theory is that individual assets have perfect substitutes. The capital asset pricing model, the arbitrage pricing theory, the Black-Scholes, and subsequent contingent claims valuation models, all rely on this notion. Various arbitrage theorems in corporate finance, such as Modigliani and Miller's (1958) Proposition I of capital structure irrelevance, also rely on perfect substitutes. Scholes's (1972) pioneering study of secondary offerings has long been considered the prime supporting evidence.

More recent studies, such as Asquith and Mullins (1986) and Masulis and Korwar (1986), provide results that could be consistent with the notion of finite price elasticities. This notion is supported by studies of the price behavior of common stock when new issues are brought to market (Barclay and Litzenberger (1988); Loderer, Sheehan, and Kadlec (1990)), by studies that relate the January effect in common stock returns to end of year tax loss selling (Rozeff (1986); Ritter (1988)) and by the analysis of the abnormal

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stock price behavior around ex-dividend days (Lakonishok and Vermaelen (1986)). Of course, the relevant question is not so much whether these elasticities are finite in a purely numerical sense, but rather whether they are large enough (in absolute terms) that they can be ignored for practical purposes. In this respect, two provocative papers that analyze the impact of new additions to the Standard and Poor's 500 Index imply that the price elasticity of demand for individual stocks could be unitary in the limit (Harris and Gurel (1986) and Shleifer (1986)).[1] A one percent decrease in the price of a particular stock would therefore lead to only a one percent increase in the quantity demanded. This would make the demand for the common stock of individual firms about as elastic as the demand for electricity for domestic use (Chapman, Tyrell, and Mount (1972)). Elasticities of this magnitude would imply that investment and financing decisions are not separable and that firms should favor the financing instruments with the most elastic demand.

Not surprisingly, the issue is far from being resolved, and opinions on the extant evidence diverge sharply. For instance, Brigham and Gapenski (1987), point out that "for years many academicians argued that the demand curve for a firm's stock is either horizontal or has an extremely slight downward slope and that signalling effects are minimal. Most corporate treasurers, on the other hand, have long felt that both effects exist for mature companies, and recent empirical studies confirm the treasurers' position" (p. 426). In contrast, Smith (1986) argues that "there is little empirical evidence in support of price pressure for widely traded stocks" (p. 9). And Brealey and Myers (1988) concur with that conclusion by arguing that "seen one stock, seen them all" (p. 296). One reason the issue is still unresolved is that few studies have been designed exclusively to test the notion that price elasticities of demand are finite.

This paper attempts to fill that void. We analyze the price discount exacted by the market when firms announce an offering of seasoned common stock, an event that seems to provide the most natural experiment to detect evidence of finite price elasticities. Unfortunately, the decision to issue stock could also convey information about future cash flows. To deal with this complication, we look first for evidence that the price decline is caused by unfavorable information about future cash flows. Then, after controlling for potential information releases, we analyze whether the residual price decline is related to possible determinants of price elasticities.

The paper differs from previous cross-sectional studies of the announcement effect of a stock offering (Asquith and Mullins (1986); Kalay and Shimrat (1987); Masulis and Korwar (1986)). Our goal is not to provide an exhaustive explanation of the price reaction, but to filter out information effects to isolate and identify the potential influence of finite price elasticities of demand.

[1] They observe that new additions to the Standard and Poor's 500 Index lead to stock price increases that are best interpreted as the result of finite price elasticities since the additions do not seem to convey new information about future cash flows.

To gauge information effects we study four variables: the price changes of nonconvertible preferred stock when firms disclose the intention to issue additional common stock, changes in analysts' earnings-per-share (EPS) forecasts, changes in actual EPS, and changes in book returns on equity. Most of our firms are regulated firms (typically utilities) since regulated firms are more likely to have preferred stock outstanding. These firms would seem to represent a most appropriate sample for our experiment since regulatory scrutiny and mandatory disclosure are likely to prevent the agglomeration of significant information asymmetries between management and capital markets. Of course, since regulated firms are likely to issue stock rather frequently, partial anticipation could also make it difficult to identify traces of finite price elasticities in their offering announcement effects.

We find little evidence that stock offering announcements by regulated firms convey adverse information about future cash flows. This is consistent with the notion that the price decline is caused by finite price elasticities of demand. Moreover, we find evidence that this reaction is related to the determinants of price elasticities postulated in an incomplete information model of capital market equilibrium developed by Merton (1987) but inconclusive evidence that it is related to other possible determinants such as liquidity and heterogeneous beliefs. On balance, the evidence suggests that issuing firms cannot treat the demand for their stock as if it were perfectly elastic.

The paper is organized as follows. Section I discusses why the price elasticity of demand for financial assets might be finite. Section II details the test design and discusses our stock offering sample. Section III estimates the announcement effect of these offerings and investigates whether the observed decline reflects the disclosure of new information about future cash flows. Section IV tests whether the decline is a function of possible determinants of finite price elasticities. Conclusions are drawn in Section V.

I. Price Elasticities of Demand for Common Stock

The demand for *individual* financial assets may have a finite price elasticity for a variety of reasons. Investors may value liquidity, that is, the ability to trade cheaply and with little delay in reaction to new information (supporting evidence is provided in Amihud and Mendelson (1986), Kamara (1989), and Warga (1990)). Even on organized exchanges, transactions are not costless. The bid-ask spread is not zero or uniform across securities, and the opportunity to trade large amounts at the market's bid or ask price is limited —the market is not uniformly broad, deep, and resilient. As frequently argued in the literature on the bid-ask spread (see, e.g., Demsetz (1968), Benston and Hagerman (1974), or Amihud and Mendelson (1986)), traders will therefore be willing to take on large positions in illiquid assets only if appropriately compensated with a lower price. If the required discount increases with the size of the position, the aggregate demand schedule for a financial asset could be downward sloping.

Demand schedules might be downward sloping for other reasons as well.

Investors may have different reservation prices for the same security because of heterogeneous information (as in Parsons and Raviv (1985)) or different interpretations of the same information. Since risk aversion limits the amount that investors with finite endowments are willing to hold, and since it takes a lower price to induce a larger number of investors with different reservation prices to buy, aggregate demand schedules could be downward sloping. Alternatively, demand schedules could have finite price elasticities because any one firm may provide investors with a type of hedging they cannot duplicate with the shares of other firms (Mayshar (1978)), because arbitrage is insufficient to counter investor sentiment (Shleifer and Summers (1989)) or because information is incomplete and risk averse investors gravitate to different sets of risky assets (Merton (1987)).

Traditional tests of the notion of finite price elasticities have consisted of attempts to find a negative relation between price and quantity changes at the time of new stock offerings. Ignoring for the moment possible information effects, if demand schedules are downward sloping, the larger the number of additional shares issued, the larger the price decline. Many studies have therefore investigated this relation.[2] Of course, researchers have been well aware of the difficulties of interpreting regressions of stock price effects on issue size in the presence of information effects. Even ignoring information effects, however, finite price elasticities do not necessarily imply a negative cross-sectional relation between relative price and quantity changes. Finite elasticities are not only consistent with cross-sectional price-quantity relations of any sign, but also with Scholes's (1972) finding of no relation at all (see Loderer and Zimmermann (1988)).[3]

[2] Although Scholes (1972) did not find any such negative relation for a secondary offering of common stock at the distribution date, Kraus and Stoll (1972) did find a negative relation as did, to some extent, Mikkelson and Partch (1985) when they examined the announcement date effect. Marsh (1979) and Hess and Frost (1982) focused instead on primary issues of seasoned equity at the issue date but rejected the hypothesis that stock price changes and issue size are negatively related. Asquith and Mullins (1986), Bhagat, Marr, and Thompson (1985), and Hess and Bhagat (1986) considered the announcement effect of primary stock issues, but only Asquith and Mullins found a negative relation between abnormal returns and issue size. Barclay and Litzenberger (1988) examined the intraday market response to announcements of new equity issues without finding a negative correlation between price change and issue size.

[3] For example, to see that they are consistent with a positive relation, assume the following linear demand function, $p = a - bQ$, where p is the stock price, Q is the number of shares, and a and b are positive parameters. Suppose firms issue the same number of new shares, $\mathrm{d}Q = 1$, where the unit of measurement is arbitrary. The relative price change of a common stock offering is $\mathrm{d}p/p$. Total differentiation of the above equation, together with the condition $\mathrm{d}Q = 1$, yields

$$\frac{\mathrm{d}p}{p} = \frac{-b\,\mathrm{d}Q}{a - bQ} = \frac{-b}{a - bQ}$$

As one can see, the relative price change is inversely related to Q, since $\mathrm{d}p/p$ is more negative at large levels of Q and less negative at low levels of Q. As the relative quantity increase $\mathrm{d}Q/Q$ is also inversely related to Q, it follows that relative price and quantity changes are positively related.

A serious complication in the analysis is that a stock offering can disclose new information correlated with offering size (Masulis (1980); Miller and Rock (1985); Krasker (1986)).[4] The consensus seems to be that stock offerings do on average convey adverse information about firm value. As evidence, researchers point out that stock offerings have a more negative announcement effect than bond issues (Eckbo (1986); Mikkelson and Partch (1986)) and that, at least in the case of industrial offerings, stock issue announcements are preceded by significant stock price increases (Asquith and Mullins (1986); Kroajczyk, Lucas, and McDonald (1989)). Both phenomena are consistent with the asymmetric information model of Myers and Majluf (1984).

II. Test Design and Sample Description

A. Test Design

To determine whether perfect substitution is a better approximation of reality than downward sloping demand curves, we examine the stock market's reaction to a stock offering announcement. If the offering is unanticipated, evidence of finite price elasticities could in principle be found at the time of the announcement (Allen and Postlewaite (1984)). We first analyze whether there is any evidence that stock offerings convey new information about future cash flows. Second, rather than testing for a negative relation between price reaction and offering size, we examine whether, after controlling for potential information effects, the market's reaction is related to possible price elasticity determinants.

We use both market and book measures of possible information releases. In particular, we use the price reaction of preferred stock. As it turns out, this choice limits our sample almost exclusively to regulated firms, which would appear to be a fortunate coincidence, since, by the nature of the regulatory process, regulated firms are less likely to be the object of significant information asymmetries. Unfortunately, regulated firms also issue stock fairly frequently, which means that any trace that finite price elasticities might leave in stock offering announcement effects could be muddled by partial anticipation.

A1. Market Measures of Information Effects

The announcement of a stock offering can convey adverse information about current or future cash flows (Miller and Rock (1985); Myers and Majluf (1984)). If prices are not always driven by market fundamentals (Shleifer and

[4] In principle, a stock offering could also lead to wealth redistribution effects among different financial claims in the firm. Galai and Masulis (1976) provide a framework to analyze this question. Wealth transfers could be related to issue size and should therefore be distinguished from the potential effects of finite price elasticities of demand. Kalay and Sangchul (1988) report evidence consistent with wealth redistribution from stockholders to bondholders in a sample of stock offerings by utilities. As we argue in a later footnote, however, there is little evidence of wealth transfers in our sample of offerings. For simplicity, we therefore ignore them.

Summers (1989)), that announcement can also convey adverse information unrelated to the worsening of any market fundamentals. For instance, issuing firms may attract favorable media attention in the pre-offer period and therfore experience stock prices systematically higher than the fundamentals would justify. Under this scenario, the announcement of a stock offering, in a manner similar to that described in Myers and Majluf (1984), could then cause noise traders to realize that managers have a lower valuation. Stock prices would decline in spite of unchanged market fundamentals.

If stock offering announcements release adverse information, they should depress the prices of all risky securities of the issuing firm, including nonconvertible preferred. Fundamentals do not necessarily have to change.[5] Conversely, if the price of nonconvertible preferred does not decline, regardless of its riskiness, there is no evidence that unfavorable information has been disclosed. We therefore study the price of nonconvertible preferred around the announcement of common stock issues to see whether in fact there is a decline. That analysis is then complemented by investigating whether the offering announcement induces changes in market fundamentals. If that is the case, then financial analysts should reduce their earnings forecasts and we should observe unexpected declines in actual post-issue earnings.

The reaction of both common and preferred stock is measured as follows. For each announcement in the sample, an event study is performed by estimating a market model for the period from 200 to 101 trading days before the announcement with rates of return defined over 2-day intervals. Since there is evidence of pre-announcement abnormal stock price behavior (Asquith and Mullins (1986)), we choose an estimation period fairly removed from the announcement date to avoid a bias in the estimates of the market model coefficients. The announcement date is the date of the first *Wall Street Journal* (*WSJ*) article to mention plans to issue additional stock. Offer announcement effects are estimated with the abnormal returns that accrue during the day preceding and the day of the *WSJ* announcement (days -1 and 0 in event time). The estimation is performed for each offering with common and nonconvertible preferred stock returns, separately. Abnormal returns for security j are defined as

$$AR_{jt} = R_{jt} - \left(a_j + b_j R_{mt}\right), \quad t = -100, \cdots, -51 \tag{1}$$

where t is a 2-day period in event time, a_j and b_j are market model parameter estimates, R_{jt} is the 2-day rate of return on security j, and R_{mt} is the 2-day rate of return on the Center for Research in Security Prices (CRSP) equally weighted index.

[5] It could be argued that dividends on nonconvertible preferred are generally contractually fixed and that therefore stock offering announcements should have a trivial impact on these securities. The evidence, however, does not support this conclusion. Kalay and Shimrat (1987) show that bonds, i.e., securities similar to preferred except for an even higher priority, are adversely affected by the announcement of common stock offerings. We will investigate whether our results depend on the riskiness of the preferred stock.

Offer announcement effects are averaged across all firms. Since returns are approximately normally distributed and cross-sectionally independent (the stock offerings in the sample occur at different times), a *Z*-statistic can be used to test for significance. Appendix A provides the necessary details.

A2. Book Measures of Information Effects

As mentioned, the abnormal return analysis is complemented by investigating whether the offering announcement induces changes in market fundamentals. The first possibility is that financial analysts lower their EPS forecasts. If these forecasts are correlated with expected future cash flows and if the announcement discloses unfavorable information about these flows, we should observe lower forecasts. Information on these forecasts is obtained from the *Institutional Brokers Estimate System* (*IBES*), compiled by Lynch, Jones and Ryan, New York. Due to lack of data for the few industrials in the sample, we concentrate on regulated firms. For each offering event, we compute

$$\Delta \mathrm{EPSF}_{T+j} = \log \frac{\mathrm{EPSF}_{T+j}}{\mathrm{EPSF}_{T-i}}, \tag{2}$$

where the subscript refers to the offer announcement month, EPSF_{T-i} is the median financial analysts' EPS forecast i months before the offer announcement date T, and EPSF_{T+j} is the analysts' revised forecast j months after the offer announcement. EPSF_{T+i} and EPSF_{T+j} refer to forecasts for the same post-offering fiscal year. $\Delta\mathrm{EPSF}_{T+j}$ is averaged across firms, and we test whether the figure we obtain is significantly less than zero. We also perform a sign-rank test to establish whether the median $\Delta\mathrm{EPSF}_{T+j}$ is significantly less than zero.

Even in the absence of adverse information, our measures of EPS forecast changes will tend to be negative simply because the offering should in principle reduce financial leverage and thereby also the risk of common stock. A lower risk level commands a lower expected return and hence lower expected EPS. To compensate for this effect, we adjust the post-offering forecasts to reflect the same number of shares as the pre-offering forecasts. To do so we multiply the post-offering forecasts by the ratio of post-offering shares divided by pre-offering shares outstanding. This new variable is denoted AΔEPSF. Significance tests of post-offering changes in adjusted EPS forecasts are the same as those outlined above for the unadjusted forecast changes. Since income available for common probably increases after the issue (presumably, the new shares are used to finance positive net present value projects or to retire debt), this adjustment will tend to inflate EPS forecasts. Together, however, adjusted and unadjusted changes in the forecasts should enable us to bound the true changes.

As an alternative investigation of unfavorable information effects, we examine earnings changes in the years surrounding the offer announcement with the method used by Healy and Palepu (1990) in their study of industrial offerings. We define standardized earnings changes for firm j in year t, ΔEPS_{jt}, as

$$\Delta\text{EPS}_{jt} = \frac{\text{EPS}_{jt} - \text{EPS}_{jt-1}}{P_j} \tag{3}$$

where EPS is firm j's COMPUSTAT annual earnings per share before extraordinary items, the time index t denotes fiscal year t, and P_j is firm j's stock price at the end of the fiscal year prior to the offering announcement year. EPS figures and stock prices are adjusted for stock splits and stock dividends.

For each issuing firm and year, we also compute a value weighted average of the standardized EPS changes for firms in the same industry. Since COMPUSTAT data are available for only a small number of industrial firms in the sample, we restrict the analysis to regulated firms and use COMPUSTAT's "utilities-composite" index for these firms. This index is reported on COMPUSTAT's aggregate tapes. To adjust for industry wide changes in EPS, we compute industry adjusted standardized EPS changes, $\text{IA}\Delta\text{EPS}_{jt}$, as the difference between the standardized EPS changes of the issuing firms and their respective industry averages. We test whether the average and median $\text{IA}\Delta\text{EPS}_t$ decline after the issue year.

The above industry adjustment procedure assumes that EPS figures of all firms in a given industry have the same correlation with the industry index. To investigate industry adjusted EPS changes without that assumption, we focus on book returns on equity (ROE) and perform a market model type regression for each firm where

$$\text{ROE}_t = \gamma_0 + \gamma_1\text{ROE}_{I,t} + \varepsilon_t, \quad t = 1, \cdots, m, \tag{4}$$

and t refers to a particular fiscal year, m is the number of annual observations available, ε_t is an independent, identically distributed error term, and $\text{ROE}_{I,t}$ is the industry's book return on equity index for fiscal year t. Again, because of sample size, analysis is restricted to regulated firms and ROE_I is measured with the utilities-composite index. ROE is defined as the ratio of income available for common (before extraordinary items) divided by common equity—COMPUSTAT items (#18 − #19)/#60. The regression is estimated using all the data available for any issuing firm on the annual 1988 COMPUSTAT tapes. For each event there are at most 19 data points ($m \leq 19$).

To test whether the disclosure of plans to issue stock is associated with lower future equity rates of return, we analyze whether the ROE observed after the issue year is lower than that predicted by equation (4). For each event, we therefore compute within sample prediction errors as

$$\text{AROE}_k = \text{ROE}_k - g_0 - g_1\text{ROE}_{I,k}, \quad k = 1, 2, \cdots, m \tag{5}$$

where g_0 and g_1 are the OLS estimators of γ_0 and γ_1 in equation (4) and k is the kth fiscal year after the offering. Prediction errors are cumulated over event time and across events and averaged. We test whether the average cumulative ROE prediction error over a certain number of years after the offering is less than zero. We also use a sign-rank test to gauge whether the median cumulative prediction error is similarly smaller than zero.

A3. Announcement Effects and Finite Price Elasticities of Demand

The preceding measures of new information can be used to isolate the components of the offering announcement stock price change potentially associated with finite price elasticities of demand. We can then test whether these residual price changes are indeed systematically related to possible determinants of demand elasticities.

Merton (1987) derives an explicit relation between price elasticity of demand and its determinants. According to the Merton model, given the risk free rate, the elasticity of the *expected return* on firm j with respect to an increase in firm size (a stock offering) is a positive function of return variance and of firm size and a negative function of the degree of "investor recognition" for stock j.[6] The intuition for these partial effects is straightforward. A higher return variance increases the compensation that risk averse investors require to hold more shares of the stock. A higher compensation is also necessary if investors have already tied up a substantial fraction of their wealth to hold the stock of a large firm. All else being equal, however, the additional compensation varies inversely with the size of the investor clientele holding stock j, because it is easier to distribute a new issue if the investor base is large.

Of course, to yield a higher post-offering expected return, the value of stock j has to fall. The size of this decline is directly related to the required increase in expected return. Consequently, the elasticity of the stock *price* with respect to offering size should be inversely related to return variance (VAR) and firm size (SIZE), and positively related to the size of the investor base (BASE). Notice that the elasticity of the stock price with respect to offering size (INVELAS) is the inverse of the price elasticity of demand as commonly defined.[7]

To test whether demand elasticities are indeed finite, we can estimate the following regression equation at the time of a stock offering announcement:

$$\begin{aligned}\text{INVELAS}_j &= \frac{\Delta P_{Cj}}{P_{Cj}} \times \frac{S_j}{\Delta S_j} \\ &= \mu_0 + \mu_1 \text{VAR}_j + \mu_2 \text{SIZE}_j + \mu_3 \text{BASE}_j + \mu_4 \text{INFO}_j + \varepsilon_{jt}, \\ & \qquad j = 1, \cdots, N \qquad (6)\end{aligned}$$

where the subscript j denotes issuing firm j, $\Delta P_C / P_C$ is the relative price change of common stock observed at the time of a stock offering announcement, $S / \Delta S$ is the inverse of the announced relative increase in the number

[6] Amihud and Mendelson (1989) have tested this model in a different context with mixed results.

[7] The Merton model assumes an inelastic supply of common stock. In our test, we also assume that the investor base does not increase at the time of the offering.

of shares of common represented by the offering, and ε_t is an independent, identically distributed error term. INFO is a measure of new information conveyed by the offering (good news corresponds to a positive value of INFO, and bad news to a negative value) and is used to filter out the confounding influence of information disclosure on the dependent variable. From the above discussion we expect μ_1 and $\mu_2 < 0$, $\mu_3 > 0$, and $\mu_4 > 0$.

Other price elasticity determinants have been suggested in the literature outside a formal theoretical framework. We investigate whether the evidence supports some of these alternative considerations. As mentioned in Section II, one can find arguments that price elasticities are related to heterogeneous reservation prices (HET) and liquidity (LIQ). To test these predictions, we measure the significance of the coefficients of proxies for HET and LIQ in a regression equation similar to equation (6). The coefficient of HET should be negative (demand functions for assets with more heterogeneous information are steeper), and that of LIQ should be positive (demand functions for more liquid assets are flatter).

B. Sample Description

The sample consists of 430 announcements of primary issues of seasoned common stock (i.e., offerings by firms of additional common stock for cash) from 1969 to 1982. The sample selection criteria are as follows:

1. The offering is recorded in an issue of the *Directory of Corporate Financing* (*DCF*) published between 1970 and 1983 by the *Investment Dealers Digest*—no offering is announced and later withdrawn;
2. The issue is reported in the *Wall Street Journal* (*WSJ*). The first mention of the issue to appear in the *WSJ* no later than the actual issue date (as reported by *DCF*) is used as the offering announcement date;
3. The issuing firm has both common and nonconvertible preferred stock outstanding at the time of the announcement, and these stocks are reported in the *Investment Statistics Laboratory* database of daily security returns and daily trading volumes for the period encompassing the *WSJ* announcement date;[8] and
4. The *WSJ* announcement does not contain information about combination offerings with other firm securities or with secondary offerings.

The 430 announcements in the sample are distributed relatively evenly over time—no single year has more than 45 or less than 21 announcements, with the exception of 1969 and 1970, which have only 1 and 12, respectively. In 11 out of 430 cases, preferred stock is not traded on the announcement date. Thus, the sample of preferred stock has only 419 announcements.

[8] Since firms that issue common stock have different types of preferred stock outstanding, we originally randomly selected the type of preferred stock to consider. That yielded 430 events with nonconvertible preferred data and 75 events with convertible preferred data. Those 75 events were later discarded.

The sample is dominated by firms in regulated industries with two-digit SIC code numbers 48 (telephone communications, telegraph, and radio and television broadcasts) and 49 (electric, gas, and sanitary services); the vast majority of firms are in industry 49. There are also a few industrial firms. Of the 430 issue announcements in the sample, 409 are by regulated firms and 21 by industrial firms. The reason for the heavy representation of regulated firms in the sample is the requirement that issuing firms have preferred stock outstanding.

As mentioned, the predominance of regulated firms could attenuate the potentially confounding influence of information disclosure in the sample, since regulatory scrutiny and mandatory disclosure could reduce the asymmetry of information between management and capital markets concerning future cash flows. Consistent with this, while industrial firms realize a cumulative abnormal return of 40 percent from 2 years until 10 days preceding the issue, utilities endure a 19 percent decline during the same period (Asquith and Mullins (1986)). There is therefore little a priori evidence to suggest information effects in our sample similar to those suspected for industrial offerings. A pre-announcement rise in stock prices is consistent with the claim that managers issue stock when it is overpriced and is perhaps the single most substantial piece of evidence in support of information effects (Korajczyk, Lucas, and McDonald (1989)).

While a sample dominated by regulated firms has the attractive characteristic that stock issue announcement effects are less contaminated by information disclosure, it also has a potential drawback, namely that regulated firms issue stock rather frequently.[9] As a consequence, the stock offering announcements of these firms are probably partially anticipated which could complicate the search for evidence of finite price elasticities.

III. The Information Content of Stock Offering Announcements

The event study results are reported in Table I which shows average and median abnormal returns and corresponding significance test statistics for both common and nonconvertible preferred stock. The sample is analyzed in aggregate and separately for regulated and industrial firms.

For the full sample, the announcement of a new offering causes a significant abnormal common stock return of -0.9% ($Z = -11.1$), and a significant proportion, 70%, of the offerings is associated with a relative price decline ($Z = 8.2$).[10] The average abnormal return is -2.8% for the industrial issuers and -0.8% for the regulated issuers (both figures are statistically significant). Asquith and Mullins (1986), Hess and Bhagat (1986), Masulis and Korwar (1986), and Mikkelson and Partch (1985) find similar results.

[9] The average number of offerings per regulated firm during our sample period is 7.0, compared with 1.3 for industrials.

[10] This Z statistic given by the ratio $(p - n \times 0.5)/(n \times 0.5 \times 0.5)^{0.5}$, where p is the number of negative returns and n the sample size. For large samples, the ratio is distributed standard normal under the hypothesis that the expected value of p is $n \times 0.5$. See de Groot (1975), p. 229.

Table I
Announcement Period Abnormal Returns

Announcement period average and median abnormal returns for common and nonconvertible preferred stock around the announcement of a primary issue of seasoned common stock. The abnormal returns are 2-day market model abnormal returns measured on the day before and the day of the offering announcement in the *Wall Street Journal*. The significance of the proportion of negative returns is tested with the statistic $Z = (p - n \times 0.5)/(n \times 0.25)^{0.5}$, where p = observed number of negative returns and n = sample size. Z is distributed unit normal for large samples under the null hypothesis that the expected value of p is $n \times 0.5$. See De Groot (1975), p. 229. The sample period is from 1969 to 1982.

Panel A. Common Stock Abnormal Returns

Sample	Sample Size	Average Abnormal Return (%)	Average Abnormal Return Z-statistic	Proportion Negative Returns (%)	Proportion Negative Returns Z-statistic
Full Sample	430	−0.93	−11.10*	69.8	8.20*
Regul. firms	409	−0.84	−10.35*	68.9	7.66*
Industrials	21	−2.82	−4.52*	85.7	3.27*

Panel B. Nonconvertible Preferred Stock Abnormal Returns

Sample	Sample Size	Average Abnormal Return (%)	Average Abnormal Return Z-statistic	Proportion Negative Returns (%)	Proportion Negative Returns Z-statistic
Full Sample	419	0.06	0.56	53.2	1.32
Regul. firms	398	0.08	0.82	52.8	1.10
Industrials	21	−0.28	−1.06	61.9	1.09

*Denotes significance at the 0.05 level (one-tail test).

By contrast, the reaction of nonconvertible preferred is not significantly different from zero. The average reaction has a value of 0.1% ($Z = 0.6$), the median (not shown) is equal to −0.03%, and there are about equal proportions of negative and positive abnormal returns ($Z = 1.3$). Insignificant average and median abnormal returns obtain for both the regulated and the industrial sample. These results provide little evidence to support the contention that primary issues of common stock by firms in our sample release unfavorable cash flow information on average. There is also little evidence to support the claim that managers' decision to issue stock leads noise traders to lower their reservation prices for the issuing firms' securities.

Of course, no change would be expected if nonconvertible preferred stock were a riskless security. Yet, when we partition the sample of regulated firms by Moody's preferred stock rating categories, we find no evidence that the riskier categories of nonconvertible preferred show a more negative announcement effect (Table II, Panel A). Preferred stock rated "ba" (the riskiest category in the sample) experiences an average abnormal return of 0.8%, compared with −0.4% for the safest category (rated "aaa").[11] None of

[11] According to *Moody's Manuals*, an issue rated "aaa" is considered to be a "top-quality preferred stock." By comparison, an issue which is rated "baa" is considered to be a "medium grade preferred stock, neither highly protected nor poorly secured." An issue which is rated "ba" is considered to have "speculative elements and its future cannot be considered well assured."

Table II

Preferred Stock Credit Ratings and Abnormal Returns

Relation between announcement abnormal return on nonconvertible preferred stock and Moody's preferred stock ratings. Regulated firms sample from 1969 to 1982. The announcement period abnormal return is defined as the 2-day market model abnormal return measured on the day before and the day of the offering announcement in the *Wall Street Journal.* The significance of the proportion of negative returns is tested with the statistic $Z = (p - n \times 0.5)/(n \times 0.25)^{0.5}$, where p = observed number of negative returns and n = sample size. Z is distributed unit normal for large samples under the null hypothesis that the expected value of p is $n \times 0.5$. See DeGroot (1975), p. 229.

Panel A. Abnormal Returns by Rating Category

Moody's Preferred Stock Rating	Sample Size	Average Abnormal Return (%)	Abnormal Return Z-statistic	Proportion Negative Returns (%)	Proportion Negative Returns Z-statistic
aaa	5	−0.43	−1.18	80.0	1.34
aa	58	0.37	1.46	46.5	−0.53
a	117	−0.03	−0.03	57.3	1.57
baa	80	0.10	0.76	47.5	−0.45
ba	22	0.78	1.46	50.0	0.00
	282				

Panel B. Analysis of Variance in Preferred Abnormal Returns

Source of Variation	Sum of Squares	Degrees of Freedom	Mean Square	F-Statistic	Significance of F
Rating	0.002	4	0.000	1.232	0.297
Residual	0.099	277	0.000		
Total	0.101	281	0.000		

these returns is significantly different from zero. There is also no evidence that risky preferred has a larger proportion of negative announcement effects than less risky preferred. An analysis of variance confirms the absence of a relation between rating and announcement effect (Panel B). Consequently, there is no evidence to suggest that stock issues release adverse information about the firm on average.[12]

For an alternative source of evidence on adverse information, we consider the changes in analysts' EPS forecasts reported in Panel A of Table III. Since *IBES* data are available for only a small number of industrial firms in the sample, we restrict the analysis to regulated firms. The *IBES* database begins in January 1976. For each firm, the database contains the median

[12] As mentioned above, there could also be wealth redistribution from common to preferred stockholders. If so, the most risky preferred stock should experience the highest benefit. There is no indication that this is occurring. In addition, if the new funds are used in scale expansion projects, the increased number of shares of common should reduce the systematic risk of the preferred (because of a decrease in financial leverage) and thereby increase the preferred's value. However, there is little evidence that the systematic risk of preferred declines. The average change in systematic risk from the pre-offer to the post-offer period is 0.001 with a t-statistic of 0.06; there is a decline in only 51% of the events.

EPS estimates by a panel for security analysts on a month by month basis. Two pairs of median forecasts are compared; the median forecasts formulated 2 months before versus 2 months after the announcement month, and the median forecasts formulated 2 months before versus 4 months after the announcement month.[13] The second pair is chosen because it may take analysts more than 2 months to revise their predictions.[14] All forecasts refer to the fiscal year after the offer announcement. On balance, the results are consistent with the above stock price findings. There is little evidence that stock offerings induce financial analysts to lower their EPS forecasts.[15]

Compare, for instance, the forecasts made in month $+4$ with those made in month -2 (data are available for only 164 events). The average unadjusted change (i.e., the change measured without taking into consideration the increase in shares outstanding) is -0.009 with a p-value of slightly more than 0.05 (one-sided t-test), and the average adjusted change is 0.094 with a p-value of 0.00. Although the p-value of the one-sided t-test for the average unadjusted change is almost 0.05, the evidence of a decline is very weak, since none of the median changes is negative. Hence, even when biasing the analysis toward finding a decline (by not adjusting for the increase in number of shares outstanding), we find little if any evidence of a significant decrease. The adjusted forecasts suggest an increase, but that could be the result of the upward bias pointed out in Section II.

The analysis can be replicated using EPS forecasts for the current fiscal year (the year of the offer announcement). The results are not shown in a separate table. In several cases the issue announcement falls very close to the end of the fiscal year, and there are therefore no post-announcement EPS forecasts for the current fiscal year. Fortunately, there are more firms with EPS forecasts for the current fiscal year than for later years. Hence, sample size is comparable to that of Panel A in Table III. As it turns out, analysts appear on average to reduce their forecasts—average EPS forecasts in month $+2$ are significantly smaller than those in month $-2 (t = -2.2)$. But not all analysts reduce their forecasts. In fact, there is an almost even split between forecast reductions on one side and unchanged forecasts or increased forecasts on the other—comparing EPS forecasts in month $+2$ with those in month -2, there are 82 reductions in a sample of 158 (sign Z-test value =

[13] EPS forecasts are compiled as of the 3rd Thursday of the month. Therefore, depending on the date of the announcement, the -2 month forecast could be an average $1\ ^1/_4$ to $2\ ^1/_4$ months prior to announcement; similarly, the $+4$ month forecast could be an average $3\ ^3/_4$ to $4\ ^3/_4$ months after the announcement. See Jain (1989).

[14] On the basis of conversions with analysts at Lynch, Jones and Ryan, Inc., Jain (1989) suggests that most analysts revise their forecasts within 4 months of a particular event.

[15] As we pointed out, our sample is dominated by public utilities. It is important to point out that the stock of public utility issuers steadily underperforms the market in the 2 years preceding the announcement (Asquith and Mullings (1986)). If the opposite were the case, given that it takes analysts some time to adjust their forecasts, the finding of unrevised post-announcement EPS forecasts could in principle reflect two offsetting influences: the good new implicit in the pre-announcement stock price performance and the bad news of the offering announcement itself. This does not seem to be a problem in a sample like ours, dominated by public utilities.

0.48); similar results obtain when comparing month +4 forecasts with month −2 forecasts (Z = 0.85). Moreover, post-announcement adjusted forecasts are higher rather than smaller than pre-announcement adjusted forecasts (t = 10.4). More importantly, the apparent downward revision in EPS forecasts is unrelated to the stock market reaction. Both Pearson and Spearman rank correlation coefficients between issue announcement effects and EPS forecast revisions for the current fiscal year are insignificantly different from zero.

We also find no evidence of a revision in long-term EPS forecasts as measured by analysts' EPS long-term growth forecasts reported on the *IBES* database. Stock offering announcements leave these forecasts unaffected—both parametric and nonparametric tests reject the hypothesis of a revision (not reported in a table). Overall, there is only weak evidence at best to suggest that stock offering announcements induce analysts to lower their EPS forecasts.[16] More importantly, there is no evidence that whatever reductions there might be are related to the stock offering announcement effects.

We also investigated whether the systematic risk of common stock changes after the offering (not reported in a table). Even if earnings are not expected to change, and increase in the systematic risk of common would decrease its value. The riskiness of common, however, does not change: the average market model beta goes from 0.41 before the offering announcement to 0.40 afterward. The change is statistically insignificant.

Hence, there is little evidence in the first panel of Table III to support the notion that stock offering announcements by regulated firms convey adverse information about future cash flows. A similar conclusion is suggested when we piece together the available evidence on industrial offerings. When studying a sample of these offerings from 1966 to 1981, Healy and Palepu (1990) report no decline in analysts' earnings forecasts on average. Jain (1989) reaches the same conclusion when investigating 281 industrial offerings in 1979–1983. Both studies find an increase in equity risk, but Jain finds no relation between offer announcement effects and equity risk increases.

Panel B of Table III shows industry adjusted EPS changes in the years surrounding the offer announcement year. Again, the analysis is limited to regulated firms because data are lacking for the industrial sample. According to our calculations, issuing firms' EPS changes always fall below the corre-

[16] For completeness, we replicated the analysis by comparing EPS forecasts (for both the current and the next fiscal year) in month −1 with those in months +2 and +4. In the text we focus on month −2 for fear that financial analysts might anticipate the stock offerings. The results of this alternative perspective are very similar to those reported in the text. The average revision in EPS forecasts is often negative and significant. Yet, the median revision is always zero (the exception being the comparison of month −1 forecasts with month +4 forecasts for the current fiscal year), and the number of negative revisions is never statistically larger than the number of nonnegative revisions (in fact, it is always significantly smaller in the case of next year forecasts). And again, adjusted changes in EPS forecasts are always positive. More importantly, changes in EPS forecasts are always uncorrelated with stock offering announcement effects (correlations computed using adjusted changes in EPS forecasts provided mixed results).

Table III

Changes in Earnings-Per-Share (EPS) Around Stock Offering Announcements

Changes in median analyst EPS forecasts, actual EPS, and excess returns on equity around stock offering announcements by regulated firms from 1969 to 1982. Panel A shows changes in the logarithm of the median analyst EPS forecasts (ΔEPSF). The forecasts are for the fiscal year following the offer announcement year. Changes are computed by comparing the logarithm of the median forecast in month $+j$ with the logarithm of the median forecast in month $-i$ relative to the offering announcement month ($i = -2$, $j = +2, +4$). AΔEPSF is ΔEPSF adjusted for the increase in the number of shares outstanding, i.e., it is computed assuming no change in the number of shares outstanding. Panel B shows changes in actual EPS as a percentage of initial equity price for the years surrounding the offer announcement year. Industry adjusted EPS changes, IAΔEPS, are the difference between EPS changes (as a percentage of initial equity price) for the issuing firms sample and the value weighted average of EPS changes (as a percentage of initial equity price) for the utilities on COMPUSTAT (as measured by COMPUSTAT "utilities-composite" index). The Panel tests the significance of IAΔEPS in different years; it also tests whether this variable becomes more negative after the offering announcement year by comparing the average (the median) value of IAΔEPS in post-announcement years with that of pre-announcement years. Panel C shows cumulative abnormal returns on equity (CAROE). Excess returns on equity are estimated error terms in a regression of firms i's annual return on equity against the contemporaneous return on a value weighted equity index for regulated firms, as measured by COMPUSTAT's "utilities-composite" index. Cumulation intervals are in years. In all panels, probability values are for one-tail tests.

Panel A. Changes in Analysts' EPS Forecasts

	Forecast month +2 vs. −2		Forecast month +4 vs. −2	
	ΔEPSF	AΔEPSF	ΔEPSF	AΔEPSF
Median	0.0	0.089	0.0	0.090
Average	−0.003	0.100	−0.009	0.094
Standard Deviation	0.058	0.098	0.072	0.109
t-test (Mean = 0), p-value	0.233	0.000	0.057	0.000
Sign Rank Test, p-value	0.337	0.000	0.105	0.000
Sample Size	164	164	164	164

Panel B. Changes in Actual Industry Adjusted EPS Changes (IAΔEPS)

Year relative to offer announcement year	Number of events	Average	t-statistic probability value	Median	Sign-rank test probability value
		Industry-adjusted EPS changes			
−5	206	−0.35%	0.004	−0.21%	0.000
−4	219	−0.35	0.006	−0.19	0.003
−3	234	−0.44	0.002	−0.37	0.000
−2	248	−0.30	0.023	−0.20	0.007
−1	257	−0.25	0.053	−0.17	0.012
0	266	−0.07	0.350	−0.18	0.048
1	266	−0.17	0.134	−0.22	0.020
2	266	−0.54	0.001	−0.22	0.001
3	266	−0.10	0.295	−0.07	0.168
4	266	−0.70	0.011	−0.30	0.005

Table III—*Continued*

Panel B. Changes in Actual Industry Adjusted EPS Changes (IA ΔEPS) (continued)

Year relative to offer announcement year	Number of events	Average	*t*-statistic probability value	Median	Sign-rank test probability value

Intertemporal comparison in industry-adjusted EPS changes (IA ΔEPS)

	year 1 vs. year −1	year 4 vs. year −1
Median	0.05%	0.07%
Average	0.08%	−0.46%
Standard Deviation	3.30%	5.28%
t-test (Mean = 0), *p*-value	0.342	0.081
Sign Rank Test, *p*-Value	0.447	0.371

Panel C. Changes in Cumulative Abnormal Returns on Equity (CAROE)

	Interval (+1, +1)	Interval (+1, +2)
Median	−0.02%	−0.24%
Average	0.13%	0.24%
Standard Deviation	2.33%	4.56%
t-test (Mean = 0), *p*-value	0.191	0.195
Sign Rank Test, *p*-value	0.364	0.102
Sample Size	266	266

sponding average industrywide changes (consistent with the negative pre-issue cumulative abnormal returns reported by Asquith and Mullins (1986) for utilities). There is no evidence, however, that this differential widens after the issue announcement. When we compare industry adjusted EPS changes 1 and 4 years after the offering announcement with those in the year before the announcement, we find little evidence to support the contention that post-issue EPS are unusually low. Industry adjusted EPS changes 1 year after the offering, for instance, are 8 basis points larger on average than during the year before the offering, and the median changes are 5 basis points larger. Neither figure is significantly different from zero. Similar conclusions follow when comparing industry adjusted EPS changes 4 years after the offering announcement with those of the year before—the average is negative and marginally significant, but the median is positive and insignificant.

Finally, Panel C of Table III reports the results of our event test using returns on equity. As one can see, the perspective changes but the results do not. There is little if any evidence of a post-offering decline in firm profitability. For instance, the average excess return on equity during the first year after the offering year is 0.1%, and the median is −0.02%. Neither figure is statistically significant.

To summarize, the results in Tables II and III lend little support to the notion that the announcement of common stock offerings discloses unfavorable information about future cash flows, particularly for regulated firms. Preferred stock is unaffected by the announcement regardless of its rating. Moreover, there is only weak evidence that investment analysts reduce their EPS forecasts; whatever revisions occur, they are unrelated to the issue announcement effects. Furthermore, EPS and accounting returns on equity do not fall. And finally, the systematic risk of common stock remains unchanged.[17] This suggests that the negative announcement effect experienced by common stock is caused by something other than adverse information, possibly finite price elasticities of demand.

IV. Stock Offering Announcement Effects and the Determinants of Finite Price Elasticities

This section provides estimates of the regression equation (6), which we repeat here for convenience.

$$\begin{aligned}\text{INVELAS}_j &= \frac{\Delta P_{Cj}}{P_{Cj}} \times \frac{S_j}{\Delta S_j} \\ &= \mu_0 + \mu_1 \text{VAR}_j + \mu_2 \text{SIZE}_j + \mu_3 \text{BASE}_j + \mu_4 \text{INFO}_j + \varepsilon_{jt}, \\ & \qquad j = 1, \cdots, N \qquad (6)\end{aligned}$$

The regression equation relates demand elasticities to their possible determinants while controlling for possible information effects. Recall that we expect μ_1 and $\mu_2 < 0$, $\mu_3 > 0$, and $\mu_4 > 0$. We also estimate a similar regression equation using proxies for heterogeneity of investors' beliefs (HET) and liquidity (LIQ) as independent variables. As mentioned before, HET should have a negative relation with INVELAS, and LIQ should have a positive relation. We first discuss the various proxies used in the estimation, then present the regression results, and finally discuss some consistency checks. Because of data constraints, we again focus on regulated firms. The few industrial issues with sufficient information yield insignificant results.

A. Proxy Variables

The proxy used for the dependent variable (INVELAS) in our regression equation (6) is the announcement period abnormal return on common stock ($\Delta P_C / P_C$) multiplied by the ratio of common shares outstanding before the new issue divided by the announced number of new shares ($S / \Delta S$). If demand schedules do not shift in reaction to possible news conveyed by a stock offering, the inverse of INVELAS measures the arc elasticity of demand for the common stock of a particular firm with respect to price. Sample descriptive statistics for this arc elasticity are provided in Panel A of Table IV. The median value is approximately -4 for regulated firms.

Our proxies for information effect (INFO) in equation (6) were discussed in Section III. For simplicity, we focus on three of the five proxies, the price change in preferred stock prices ($\Delta P_P / P_P$), the adjusted change in analysts' EPS forecasts (AΔEPSF), and the cumulative abnormal return on equity

[17] Sample sizes vary significantly from one analysis to the other depending on data availability.

(CAROE). The conclusions do not change when we use versions of the two remaining measures, ΔEPSF and IAΔEPS. Panel B of Table IV reports the Pearson correlation coefficients between pairs of these proxy variables. For completeness, we also include the announcement period abnormal return on common stock ($\Delta P_C/P_C$). None of the information proxies are significantly related to $\Delta P_C/P_C$, which corroborates the evidence that the price change of common stock of regulated firms is unrelated to new earnings information. By contrast, changes in preferred stock prices ($\Delta P_P/P_P$) are related to unexpected changes in returns on equity. This indicates that stock issues could release news about the firm, but apparently only news concerning the value of securities with a higher priority than common. The *average* price of preferred, of course, is unaffected by the offering announcement (Table I). Table IV also shows no significant correlation between changes in EPS forecasts and other information proxies. The results are the same when computing Spearman correlation coefficients, except that changes in adjusted EPS forecasts correlate positively and significantly with the cumulative abnormal return on equity.

VAR in regression equation (6) is measured by the daily common stock return variance over event days -100 to $+100$. BASE is the average daily trading volume in a particular common stock from event day -100 to $+100$, divided by the number of common shares outstanding before the offering. This proxy is denoted by VOL. And the proxy for SIZE is the natural logarithm of the pre-offer number of shares of common times the share price 2 days before the offer announcement, deflated by the year end value of the Consumer Price Index.

The following variables are used as proxies for liquidity (LIQ). Traditionally, liquidity had been hypothesized as being positively related to trading volume (see, e.g., Fisher (1959), Demsetz (1968), Benston and Hagerman (1974), and Copelend (1979)). In keeping with that tradition, LIQ can be measured by the average trading volume in a particular common stock, i.e., the same variable used to measure a given security's investor base. This proxy is also consistent with the analysis in Lippman and McCall (1986). If firms with more liquid stocks face flatter demand curves, the coefficient of trading volume in a cross-sectional regression equation such as (6) should be positive.

Liquidity can also be measured with Stoll and Whaley's (1983) relative bid-ask spread measure (B–A), computed as the average of the relative spreads existing at the beginning and the end of the year.[18] Data on this variable for the period 1973 to 1979 were kindly provided to us by H. Stoll. Assuming finite price elasticities, less liquid securities should be characterized by larger relative bid-ask spreads, which implies that the coefficient of B–A in a regression such as (6) should be negative.

The following proxies can be used to measure heterogeneity of investors' beliefs (HET). Suppose, for simplicity, that the market for a particular stock

[18] The relative spread is the ratio (ask price − bid price)/[(ask price + bid price)/2].

Table IV

Regression Variable Descriptive Statistics

Descriptive statistics on the measure of arc elasticity of demand for common stock, information disclosure proxies, and proxies for liquidity and heterogeneous beliefs for the regulated firms sample from 1969 to 1982. The arc elasticity measure is defined as the inverse of the announcement period abnormal return multiplied by the ratio of shares outstanding divided by announced new shares to be issued. Panel A presents descriptive statistics for the arc elasticity of demand measure. Panel B shows Pearson correlation coefficients between pairs of proxies for information effects. Numbers shown are correlation coefficients, probability values (in parentheses), and sample sizes. Panel C lists descriptive statistics for liquidity (LIQ) and heterogeneous beliefs (HET) proxies. The variable in Panels B and C are defined as follows:

$\Delta P_c / P_c$ = Common stock's announcement period abnormal return;
$\Delta P_p / P_p$ = Preferred stock's announcement period abnormal return;
AΔEPSF = Change in the logarithm of the median analysts' earnings-per-share (EPS) forecast, adjusted for the change in the number of shares outstanding. The change is computed by comparing the logarithms of the median EPS forecasts in month $+4$ after offering announcement with those in month -2 before offering announcement;
CAROE = Abnormal return on equity during the fiscal year following the offering announcement;
VOL = Average daily trading volume divided by the number of common shares outstanding prior to issue announcement;
B-A = Average relative bid-ask spread;
VAR = Daily common stock return variance;
CV = Coefficient of variation of security analysts' EPS forecasts.

Panel A. Arc Elasticity of Demand Measure

Average elasticity	−11.12
Median elasticity	−4.31
Minimum elasticity	−1842.11
Maximum elasticity	481.52
Number of observations	409

Panel B. Pearson Correlation Coefficients Between Information Effects Proxies

	$\Delta P_p / P_p$	AΔEPSF	CAROE
$\Delta P_C / P_C$	0.052	0.082	−0.023
	(0.300)	(0.295)	(0.704)
	398	164	264
$\Delta P_p / P_p$		0.059	0.201
		(0.456)	(0.001)
		161	256
AΔEPSF			0.135
			(0.128)
			128

Panel C. Liquidity and Heterogeneous Beliefs Proxies

Proxies for Liquidity (LIQ)		
VOL	Average	0.0017
	Median	0.0016
	Observations	411

Table IV—*Continued*

Panel C. Liquidity and Heterogeneous Beliefs Proxies (continued)		
B–A	Average	0.0110
	Median	0.0101
	Observations	254
Proxies for Heterogeneous Beliefs (HET)		
VAR	Average	0.0003
	Median	0.0003
	Observations	404
CV	Average	0.0467
	Median	0.0403
	Observations	206

was maintained by a specialist on an organized exchange. If investors' beliefs differed widely, the specialists' book would reveal a wide dispersion of buy and sell orders. All else being equal, a certain market order to buy or sell an amount exceeding what was available at the ask or bid price would therefore cause a larger price change than if buy and sell orders were less widely dispersed. A larger divergence in investors' beliefs should therefore lead to a greater stock return variability. The reasoning in French and Roll's (1986) study of trading- and nontrading-period return variances is implicitly the same.[19] Stock return variability (VAR) is therefore a possible proxy for HET. Ritter (1988) uses a similar measure of heterogeneity in expectations. VAR should have a negative coefficient estimate in regression equation (6).

The coefficient of variation in analysts' EPS forecasts (CV) provides an alternative proxy for HET. For each event, we compute CV by calculating the ratio of the standard deviation of analysts' EPS estimates for the issuing firm divided by the average EPS estimate for this same firm in each of the 12 months prior to the stock offering announcement month. The variable CV is the average of these monthly coefficients of variation. As with VAR, if demand curves are downward sloping, the coefficient of CV in regression equation (6) should be negative—more heterogeneous beliefs imply steeper demand curves.[20]

Panel C of Table IV reports descriptive statistics for all the proxies for liquidity and heterogeneous beliefs. We also computed correlation coefficient estimates between proxy pairs. Several correlation coefficients are sig-

[19] If investors' disagreement generates trading, the heterogeneity of investors' beliefs could also be measured by our liquidity measure VOL, in which case VOL should have a negative coefficient in the regression. Hence, a priori, there is some ambiguity about whether VOL measures liquidity or heterogeneity of beliefs. The purpose here, however, is not so much to identify the differential importance of these two determinants of demand elasticities as to establish whether either determinant affects the announcement effect of a stock offering. Finding such an effect would contradict the notion of perfectly elastic demand schedules.

[20] As another possible proxy for HET, we also used the average number of shares of a particular common stock sold short. It never has a significant coefficient in the regressions.

nificant, but none is large enough to raise excessive concerns about multicollinearity.

B. Results

Regression estimates of equation (6) are shown in Table V. In the first column, we estimate equation (6) directly. In columns (2)-(5), we estimate regressions similar to equation (6) with proxies for liquidity and heterogeneity of investors' beliefs. Those regressions are not intended to be a test of Merton's hypotheses but rather a test of the general notion that liquidity (LIQ) and heterogeneous investor beliefs (HET) are determinants of price elasticities of demand. Consider the results in the first column. Consistent with the evidence reported in Section III, the proxy for information disclosure ($\Delta P_P/P_P$) does not have any explanatory power in this or in any of the regression models. By contrast, all the determinants of demand elasticities have a significant coefficient with the sign predicted by Merton (1987): return variance (VAR) and firm size (SIZE) have a negative coefficient ($t = -2.7$ and $t = -6.9$) and relative trading volume (VOL) a positive coefficient ($t = 3.7$). The variable with the most explanatory power is firm size; this relation suggests that investors require a higher expected return to hold more of an already large firm.

Of course, these results also support the claim that liquidity and heterogeneous beliefs are determining factors of demand elasticities. More liquid stocks (as measured by VOL) experience a smaller price decline, while stocks with more heterogeneous information (as proxied by VAR) experience larger price declines. In the second regression (column (2)), we investigate the significance of liquidity (LIQ) and heterogeneous beliefs (HET) further and substitute the coefficient of variation of analysts' EPS forecasts (CV) for VAR as a proxy for HET. Since the corresponding data are not available for the full sample, the number of observations declines from 398 to 201. With this regression specification, the coefficient of trading volume (VOL), the liquidity proxy, maintains a positive and significant coefficient ($t = 2.0$) while the coefficient of the heterogeneous beliefs proxy (CV) and that of firm size (SIZE) lose their significance ($t = 0.8$ and $t = -0.3$, respectively).

In the third regression specification, we substitute the relative bid-ask spread (B-A) for trading volume as a measure of liquidity, while using VAR as a proxy for heterogeneous beliefs. Availability of B-A measures restricts the sample size to 245. With this specification, return variance and firm size have a significance coefficient with the correct sign ($t = -3.7$ and $t = -6.8$), while the relative bid-ask spread variable has a coefficient with the predicted, but not significant, negative sign ($t = -0.9$).

In columns (4) and (5), we replace the abnormal return on preferred stock ($\Delta P_P/P_P$) with alternative measures of information disclosure and use the regression model of column (1). The coefficients on these alternative information proxies are not significantly positive either, which further supports the notion that regulated firm stock offerings do not reveal earnings information

Table V

Cross-Sectional Regressions Using the Inverse of the Measured Price Elasticity of Demand as the Dependent Variable

The table reports the estimation results of two models:

Model 1: $\text{INVELAS}_j = \mu_0 + \mu_1\text{VAR}_j + \mu_2\text{SIZE}_j + \mu_3\text{BASE}_j + \mu_4\text{INFO}_j + \varepsilon_{jt}$
Model 2: $\text{INVELAS}_j = \mu_0 + \mu_1\text{HET}_j + \mu_2\text{SIZE}_J + \mu_3\text{LIQ}_j + \mu_4\text{INFO}_j + \varepsilon_{jt}$.

Estimation results of model 1 are reported in column 1, and estimation results of model 2 are reported in columns 1–5. Regulated firms sample from 1969 to 1982. Variables and proxies are defined as follows:

—The dependent variable in all the regressions is given by INVELAS = $\Delta P_c/P_c(S_c/\Delta S_c)$, where

$\Delta P_c/P_c$ = Common stock's announcement period abnormal return;
S_c = Number of shares of common stock outstanding before the announcement;
ΔS_c = Announced number of shares of common stock to be issued;

—The proxy for VAR is the daily common stock return variance;
—The proxy for SIZE (firm size) is the natural logarithm of the pre-offer number of shares of common times the share price 2 days before the offer announcement, deflated by the year end value of the Consumer Price Index;
—The proxy for BASE (the investor base) is VOL, the average daily common stock trading volume divided by the number of common shares outstanding before the announcement;
—Proxies for INFO (information effects) are $\Delta P_p/P_p$, AΔEPSF, and CAROE, where

$\Delta P_p/P_p$ = Preferred stock's announcement period abnormal return;
AΔEPSF = Change in the logarithm of the median analysts' earning-per-share (EPS) forecast, adjusted for the change in the number of shares outstanding. The change is computed by comparing the logarithms of the median EPS forecasts in month +4 after offering announcement with those in month −2 before offering announcement;
CAROE = Abnormal return on equity for the fiscal year following the offering announcement.

—Proxies for LIQ (liquidity) are VOL (already defined) and B–A (the average relative bid-ask spread);
—Proxies for HET (heterogeneity of investors' beliefs) are VAR (already defined) and CV (the coefficient of variation of security analysts' EPS forecasts).

Independent Variables		(1)	(2)	(3)	(4)	(5)
Intercept	coefficient	1.747	−0.027	2.122	0.201	2.164
	t-statistic	6.239	−0.050	6.477	0.337	7.051
VAR	coefficient	−153.015		−286.115	−109.411	−126.540
	t-statistic	−2.727*		−3.720*	−0.552	−1.922*
SIZE	coefficient	−0.098	−0.008	−0.110	−0.016	−0.119
	t-statistic	−6.888*	−0.290	−6.839*	−0.545	−7.702*
VOL	coefficient	53.966	38.357		39.930	51.128
	t-statistic	3.650*	2.003*		1.797*	2.712*
B–A	coefficient			−2.905		
	t-statistic			−0.886		
CV	coefficient		0.706			
	t-statistic		0.752			
$\Delta P_p/P_p$	coefficient	0.239	0.588	0.719		
	t-statistic	0.405	0.657	1.043		
AΔEPSF	coefficient				0.080	
	t-statistic				0.370	

Table V—*Continued*

Independent Variables		(1)	(2)	(3)	(4)	(5)
CAROE	coefficient					−1.157
	t-statistic					−1.752*
Adjusted R^2		0.138	0.011	0.194	−0.003	0.204
F-statistic		16.841	1.576	15.690	0.874	17.616
p-value		0.000	0.182	0.000	0.481	0.000
Sample Size		398	201	245	163	260

*Denotes significance at the 0.05 level (one-tail test). There is no prediction concerning the intercept.

that the market can use to reassess its valuation of common stock. Specifically, when information releases are approximated with analysts' adjusted EPS forecast changes (column (4)), the only (marginally) significant coefficient is the one associated with trading volume ($t = 1.8$).[21] These results are not the consequence of the way we measure information because we obtain the same coefficients when we go back to the original specification (1) (with $\Delta P_P / P_P$ as the information proxy) and reexamine the 163 observations available in column (4). These observations might be systematically different from the rest of the sample. As it turns out, this subsample of offering events has an average announcement effect smaller in absolute value than that observed for the full sample (−0.5% versus −0.8%). Hence, our conjecture of sample bias may have some merit.

When information disclosure is measured with the unexpected change in ROE (column 5), the results are similar to those in column (1). As in column (1), return variance ($t = -1.9$), trading volume ($t = 2.7$) and firm size ($t = -7.7$) are significant with the sign predicted by finite price elasticities of demand. The coefficient of CAROE is negative and marginally significant, contrary to what information effects would predict.

Overall, the regressions do not seem to explain much of the variation of the dependent variable, although that could be due to the unavoidable noise in daily stock returns.[22] If we look at statistical significance, some of the evidence supports the notion of finite price elasticities. The predictions of Merton (1987), for instance, are supported by column (1) in the table. We do not find, however, convincing evidence that proxies for liquidity and heterogeneous beliefs are determinants of demand elasticities: some proxies (VOL and VAR) have the predicted coefficients, whereas others (B-A and CV) do not.[23] The insignificant coefficients associated with the relative bid-ask

[21] The same results obtain with the unadjusted measure of EPS forecast changes.

[22] We reestimated the models in columns (2) and (3) with the alternative measures of information and obtained results with similar implications as those reported in Table V. The coefficient of VOL, however, was not significant.

[23] We also tried Roll's (1984) implicit measure of the bid-ask spread as an additional proxy for liquidity in the regression specification of column (1). As it turns out, this variable has little if any explanatory power.

spread (B–A) and the coefficient of variation in analysts' EPS forecasts (CV) are puzzling since these two variables seem to be good proxies for liquidity and heterogeneous beliefs. One possibility is that the relative bid-ask spread is not an appropriate gauge of liquidity in the case of large transactions, as it fails to measure market depth appropriately. The evidence we do find does not seem to be general either, for it appears to be driven by a subset of the events, namely, those not included in the sample of 163 investigated in column (4) of the table.

As we mentioned, regression estimates for the subsample of industrial firms are inconclusive (these results are not reported in a separate table). In the case of specification (1), all variables have insignificant coefficients, though VOL and VAR have the correct sign. There are only 21 events with all the necessary data. Among other reasons, insignificant coefficients could obtain because of the small sample size or because cash flow information effects are more pronounced in the case of unregulated firms, which would imply a regressor $\Delta P_P / P_P$ more likely to be correlated with the error term. For the other regression specifications, available data were further restricted to roughly 12 events each. Therefore, no conclusion is drawn from these regression results.

C. Consistency Checks

To shed more light on what exactly is being measured by trading volume, return variance, and firm size, we perform four consistency checks. For simplicity, we focus on the specification in column (1), which appears to provide the strongest support for the notion of finite price elasticities. Since the regression residuals appear from their Student t-test ranges to be leptokurtic, our first check (not reported in a separate table) is to test the significance of the coefficients using the consistent estimator proposed by White (1980) for the covariance matrix of the regression coefficients in the presence of an unknown form of heteroskedastic residuals. As it turns out, this alternative estimation technique does not yield substantially different results.

As a second check, we investigate whether the results depend on how we measure the dependent variable. As an alternative, we use the announcement period abnormal return as the dependent variable. Since that abnormal return could be related to the size of the issue, we include offering size as an independent variable. The results are reported in the second column of Table VI (for comparison, the first column of the table replicates the first column of Table V). Even with this different specification, the three proxies for the determinants of demand elasticities maintain their sign and significance, although at a lower level of confidence.

Third, we replace the dependent variable altogether. Instead of the inverse of what could be the price elasticity of demand, we choose our nonmarket proxies for information effects. If demand elasticities are the only factor at work, and if VAR, VOL, and SIZE really represent determinants of demand

Table VI

Cross-Sectional Regressions With Alternative Dependent Variables

The table reports the estimation results of the model:

$$Y_j = \mu_0 + \mu_1 VAR_j + \mu_2 SIZE_j + \mu_3 VOL_j + \mu_4 \Delta P_p / P_{pj} + \varepsilon_{jt}$$

The sample is the regulated firms sample from 1969 to 1982. Four different variables are used as dependent variables:

—INVELAS = $\Delta P_c / P_c (S_c / \Delta S_c)$, where

$\Delta P_c / P_c$ = Common stock's announcement period abnormal return;
S_c = Number of shares of common stock outstanding before the announcement;
ΔS_c = Announced number of shares of common stock to be issued;
—$\Delta P_c / P_c$ = Common stock's announcement period abnormal return;
—AΔEPSF = Change in the logarithm of the median analysts' earning-per-share (EPS) forecast, adjusted for the change in the number of shares outstanding. The change is computed by comparing the logarithms of the median EPS forecasts in month +4 after offering announcement with those in month −2 before offering announcement;
—CAROE = Abnormal return on equity during the fiscal year following the offering announcement;

The independent variables are measured as follows:

—VAR = Daily common stock return variance;
—SIZE = Natural logarithm of the CPI deflated market value of equity;
—VOL = Average daily common stock trading volume divided by the number of common shares outstanding before the announcement;
—$\Delta P_p / P_p$ = Preferred stock's announcement period abnormal return;
—$\Delta S_c / S_c$ = Announced relative increase in the number of common shares outstanding.

Independent Variables		Dependent Variable: INVELAS	$\Delta P_c / P_c$	AΔEPSF	CAROE
Intercept	coefficient	1.747	0.057	0.593	0.044
	t-statistic	6.238	2.138	2.780	1.511
VAR	coefficient	−153.015	−12.102	−20.338	17.312
	t-statistic	−2.727*	−2.942*	−0.280	2.826
SIZE	coefficient	−0.098	−0.003	−0.026	−0.003
	t-statistic	−6.888*	−2.758*	−2.394*	−1.950*
VOL	coefficient	53.966	2.252	2.224	4.159
	t-statistic	3.650*	2.069*	0.274	2.361*
$\Delta P_p / P_p$	coefficient	0.239	0.028		
	t-statistic	0.405	0.637		
$\Delta S_c / S_c$	coefficient		0.018		
	t-statistic		0.631		
Adjusted R^2		0.138	0.051	0.019	0.067
F-statistic		16.841	5.224	2.059	7.193
p-value		0.000	0.000	0.106	0.000
Sample Size		398	398	163	261

*Denotes significance at the 0.05 level (one-tail test). There is no prediction concerning the intercept.

elasticities (as opposed to residual information effects not captured by the change in the price of preferred), these variables should not be related to book measures of profitability. The results, reported in the third and fourth column of Table VI, seem to weaken our evidence.

Specifically, volume has an insignificant coefficient when the dependent variable is the analysts' adjusted EPS forecast revision (third column), but a significantly positive coefficient (just as in the second column) when the dependent variable is the abnormal change in return on equity (fourth column). While this result cannot be explained with information effects (it is not clear how volume could correlate with new information and not affect analysts' EPS forecasts), it cannot be explained with finite demand elasticities either. Finite price elasticities also fail to explain why firm size is related to both book measures of profitability. The only result that supports finite price elasticities is that return variance has either an insignificant coefficient (third column) or a positive and significant coefficient (fourth column). As one may remember, return variance has a negative and significant coefficient in almost all of our previous regressions. On balance, these findings suggests that the relation between INVELAS and trading volume and firm size could be caused by factors other than finite price elasticities. Alternatively, it could also be the case that trading volume and firm size are proxies for finite price elasticities in the regressions with INVELAS and proxies for other factors in the regressions with book measures of profitability. By contrast, the role of variance seems to be tied to finite price elasticities.

Before we can settle on this conclusion, however, we have to more carefully consider the possibility that return variance really captures aspects of systematic risk that our abnormal return measurement approach is unable to filter out. Our last consistency check is therefore to explore this scenario (there is no separate table for this analysis). We assume an imaginary event on days +201 and +202 after the offer announcement and compute abnormal returns for those 2 days. This imaginary event is chosen far from the stock offering announcement date to reduce the likelihood of issue-related price changes. With these abnormal returns as the dependent variable, we estimate the same cross-sectional regression as in the second column of the table. If return variance has a significant coefficient in our regressions simply because we fail to properly account for risk premia, then it should have a significant coefficient in this additional regression as well. But that is not what we find: return variance does not enter the regression with a significant coefficient (and neither do any of the other variables). There is consequently no evidence that return variance correlates with residual risk effects.[24]

[24] We also inquired whether trading volume measures the liquidity of the individual securities analyzed or the overall liquidity of the market at the time of a stock offering. Market wide liquidity (MVOL) is measured by the 10-day average trading volume reported 2 days before the offer announcement date by *Standard and Poor's Daily Stock Price Record* for the exchange on which the firm in question trades. When both VOL and MVOL are included, only VOL has a positive and significant coefficient, whereas the coefficient of MVOL is not significantly different from zero.

IV. Conclusions

The purpose of this paper is to find evidence of finite price elasticities of demand for the common stock of individual corporations. We investigate the changes in the price of common stock experienced by firms that announce a common stock offering. Most of the analysis deals with regulated firms, since these firms typically issue preferred stock, and preferred stock price changes are among the variables used to investigate new information released by the offering announcement. Regulated firms provide the sample of offering events least likely to be contaminated by information disclosure. Unfortunately, since regulated firms are also relatively frequent issuers of common, their offering announcements do not come as complete surprises. Partial anticipation could therefore complicate the search for evidence of finite price elasticities.

According to our results, the issue announcement by a regulated firm causes a price decline that cannot be explained by a simultaneous release of adverse information about firm value or about future cash flows (if prices are not always driven by fundamentals, these two possibilities are not one and the same). Neither the price of nonconvertible preferred nor analysts' EPS forecasts nor accounting measures of profitability (EPS and return on equity) are adversely affected in a significant way by this announcement. The decline cannot be explained by risk increases either, since common stock risk is unchanged. As a consequence, the observed decline in the price of common stock could reflect finite price elasticities of demand. Interestingly, similar studies of industrial stock offerings also fail to link offer-announcement price declines to unfavorable firm-specific cash flow information.

We attempt to relate the price decline observed on the announcement of the stock offering to the possible determinants of demand elasticities. We find results that are consistent with a model of finite price elasticities proposed by Merton (1987) but inconclusive evidence of an empirical tie between announcement price changes and a set of possible proxies for liquidity and heterogeneous beliefs. Overall, our main evidence is similar in nature to that uncovered by Shleifer (1986) and Harris and Gurel (1986). They observe that new additions to the Standard and Poor's 500 index lead to stock price increases that are best interpreted as the result of finite price elasticities, since the additions do not seem to convey new information about future cash flows. Here, we observe that stock offerings by regulated firms depress stock prices, but there is little evidence that this reaction is caused by the disclosure of unfavorable information. Hence, the most likely interpretation of these price declines is finite price elasticities. Our results, however, do not carry us significantly beyond that conclusion. Our attempts to relate announcement effects to determinants of price elasticities are inconclusive. One possibility is that the market's reaction to a stock offering announcement provides a noisy measure of demand elasticities. For instance, if the announcement is partially anticipated (and there is a strong presumption that is indeed the case in our sample of regulated firms), then the announcement effect reflects only a fraction of the decline that the price elasticity of demand brings about. This could make it difficult to find a cross-sectional relation between the announcement effect and elasticity determinants. The estimation problem would be compounded if elasticities changed over time, as suggested by evidence of time-varying liquidity premiums (Warga (1990)).

On balance, the evidence suggests that neither information effects nor finite price elasticities provide a fully convincing account of stock offering announcement effects by regulated firms. Finite price elasticities, however, provide the better explanation. First, there is only weak evidence that stock offerings by regulated firms disclose adverse information. This finding seems to exclude adverse information as the major cause of negative issue announcement effects. Second, at the very least, the evidence cannot exclude the presence of finite price elasticities, and in some cases it even suggests it. Of course, more work is needed to substantiate this conjecture.

Appendix A

This appendix describes the significance tests used in the event study. The abnormal return of security j for the 2 day event period t, AR_{jt}, is standardized by its estimated standard error to form a standardized abnormal return

$$SAR_{jt} = AR_{jt}/S_{jt}$$

where

$$S_{jt} = S_j\sqrt{1 + \frac{1}{T_j} + \frac{(R_{mt} - AR_m)^2}{\sum_{i=1}^{T_j}(R_{mi} - AR_m)^2}}$$

T_j is the number of 2-day returns used to estimate the market model parameters for security j, AR_m is the average 2-day return to the market index over the estimation period, R_{mi} is the return on the market index for the 2-day period i, and S_j is the square root of the estimated residual variance from the market model for security j. If the AR_{jt}'s are normally distributed, the SAR_{jt}'s are distributed Student's t and approximately unit normal. To test the statistical significance of the average abnormal return for event period t, we use $Z_t = W_t/(N)^{0.5}$, where

$$W_t = \sum_{j=1}^{N} \mathrm{SAR}_{jt},$$

and N is the number of securities in the sample. If the SAR_{jt}'s are independent with an expected value of zero, Z_t will be distributed approximately unit normal.

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