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Issuer surplus and the partial adjustment of IPO prices to public information ☆

Roger M. Edelen[a,*], Gregory B. Kadlec[b]

[a]*Mellon Capital Management, Corp., San Francisco, CA 94105, USA*
[b]*Pamplin College of Business, Virginia Tech, Blacksburg, VA 24060, USA*

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Abstract

This study develops a model in which rational issuers maximize the expected surplus from going public by choosing an offer price that weighs the benefit of higher proceeds if the offer is completed against the cost of foregone surplus if the offer fails. Increases in the market valuation of comparable firms during the waiting period imply higher surplus associated with going public; issuers respond with a partial revision in the offer price to elevate the probability of completion. The model offers insights into many facts associated with initial public offering pricing, including partial adjustment to market returns, the inverse relation between withdrawal and market returns, the asymmetric price adjustment to up versus down market returns, hot-issue markets, and unconditional underpricing.

JEL: G14; G24; G32

Keywords: Offering; IPO; Issuer; Underpricing; Withdrawal; Partial adjustment; Reputation

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*Corresponding author. Fax: 415 9752106.
E-mail address: rogere@mcm.com (R.M. Edelen).

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1. Introduction

In the typical initial public offering (IPO) an issuer retains an investment bank to act as its agent in pricing and selling shares. In this process the offering is filed, an initial price range is reported, and then the actual offer price is set after a waiting period of several weeks. Studies dating back to Logue (1973) show that the degree to which IPO shares are underpriced is positively related to the market return during the waiting period, which implies that the offer price only partially incorporates changes in market valuation during the waiting period (see Hanley, 1993). While the investment bank technically sets the offer price, the issuer reasonably has input over the price and certainly has final say over whether the offering is made at a given price. Thus, the key economic consideration regarding the partial incorporation of market returns into offer prices lies with the issuer's decision to go forth with such a price.

For the most part, the literature does not address this curious aspect of IPO pricing, but instead treats the empirical regularity as a necessary, ad hoc control. One exception is Loughran and Ritter (2002), who offer a theoretical justification for issuers' decision to condone such a pricing policy. They argue that prospect theory explains the partial adjustment phenomenon.[1] Prospect theory asserts that economic agents' preference for wealth depends on whether it is recently acquired (relatively low utility) or long-held (relatively high utility). In the context of IPOs, prospect theory asserts that issuers "bargain hard" when market valuations fall and are "pushovers" when market valuations rise. This study offers an alternative explanation in which issuers partially adjust the offer price in response to changes in market valuations during the waiting period so as to maximize their expected surplus from going public. Thus, our study distinguishes itself from Loughran and Ritter (2002) in that we use standard preferences to model the IPO pricing decision. (Derrien, 2003, also develops a behavioral model to explain the partial adjustment phenomenon, while Ljungqvist and Wilhelm, 2004, provide empirical support for the prospect theory explanation.)

When a firm chooses to go public, revealed preference suggests that the issuer expects some surplus associated with public versus private ownership. A rational issuer follows a pricing policy that maximizes the expected value of that surplus. That is, pricing decisions are made considering both the proceeds conditional on deal success and the probability of deal success. If the issuer's surplus at the expected public valuation of the firm is large, the rational issuer wants the offering priced conservatively (i.e., underpriced more) because the marginal benefit to increasing the probability of completion is high. Conversely, if the issuer's surplus at the expected public valuation of the firm is small, the issuer wants the offering priced aggressively

[1] A number of studies examine the partial adjustment of offer prices to private information learned by the issuer or bank during the waiting period (e.g., Benveniste and Spindt, 1989; Hanley, 1993). However, these book building models, developed to account for partial adjustment to private information, do not address the partial adjustment to public information.

(i.e., underpriced less) because the marginal cost of increasing the probability of failure is low.

Partial adjustment of IPO prices to market returns during the waiting period follows from this policy. The most common means for determining the expected public valuation of the IPO firm is to use the market value of publicly traded comparable firms. To the extent that the issuer's surplus from going public is positively related to changes in comparable firm value, our model predicts that a rational issuer only partially revises the offer price in response to changes in comparable firm value during the waiting period.

The model rests on two key assumptions. First, an issuer who pushes for a more aggressive price faces a greater likelihood that withdrawal is forced by the investment bank. While this could be motivated in several ways, we focus on the bank's reputation cost. If unconstrained by the bank, a rational issuer would direct the bank to make an infinite number of solicitations to investors, beginning at a very high price and lowering infinitesimally until just meeting either the maximum price that investors are willing to pay or the minimum price that the issuer is willing to accept. The bank, however, does not agree to subject investors to endless rounds of solicitation because of reputation costs. If a solicitation fails at one price, the bank requires that the next solicitation occurs at a price sufficiently lower to maintain its reputation. When this constraint is below the issuer's reservation price the offer is withdrawn.[2] Second, the issuer's surplus from going public is increasing in comparable-firm value. This would occur, for example, if the value of the firm as a private entity increases less than the value of the firm as a public entity because it is unable to exploit positive net present value projects because of capital constraints. (This assumption is also consistent with diversification as a motive for going public. As the value of the firm increases the marginal utility from diversification increases, thus, implying greater surplus from going public.)

In addition to explaining partial adjustment to waiting period market returns, the model offers insights into several other previously explained empirical regularities regarding IPOs. First, on average IPOs are underpriced (Ibbotson, 1975). To the extent that conducting an IPO is costly, a private firm chooses to go public only when the ex ante surplus at the expected offer price is sufficiently large to cover the expected cost. After filing, some of these costs are sunk, hence, the surplus at the expected offer price is positive. According to the model, with positive surplus the issuer underprices to increase the probability of completion.

Second, IPO underpricing is correlated both across offerings and over time, the so-called hot issue market (Ibbotson and Jaffe, 1975; Ritter, 1984). Both the cross-sectional and time-series correlations follow from the model's central prediction that offer prices are revised only partially in response to changes in comparable firm

[2]This type of constraint could also arise from a bank's search costs. The higher the offer price relative to expected public value, the more costly it is for the bank to find sufficient demand (Baron and Holmstrom, 1980). Alternatively, it could be motivated by a firm-commitment offering wherein the investment bank seeks to maintain low underwriting risk.

value.[3] The cross-sectional correlation in underpricing arises because IPO firms with overlapping waiting periods partially adjust in response to a common factor. The time-series correlation arises for reasons similar to the autocorrelation in portfolio returns caused by nonsynchronous trading. Suppose that there is a large market return in week t. During this week some IPO firms are near the end of their waiting period, while others are only beginning the process. The market return in week t affects underpricing for both issuers. However, the effect on underpricing is observed asynchronously, because the two issuers complete their offerings at different times. Thus, the time series of IPO underpricing exhibits serial correlation.

Third, the frequency of IPO withdrawal is inversely related to waiting-period market returns (Busaba et al., 2001; Benveniste et al., 2003). This result follows from the model's prediction that the aggressiveness with which an offer is priced is inversely related to changes in comparable firm value during the waiting period. We provide additional evidence on the relation between withdrawal and waiting period market returns that is consistent with our model. In particular, the inverse relation between withdrawal and comparable firm returns is driven by an increase in withdrawal in down markets. In up markets, essentially no relation exists between withdrawal and comparable firm returns.

Finally, the apparent reliance on underpricing instead of gross spreads as a choice variable in initial public offerings (Chen and Ritter, 2000; Hansen, 2001) arises naturally within the framework of our model. Raising the gross spread puts more money into the investment bank's hands, which perhaps is not the most efficient mechanism to improve the probability of completion. Suppose, for example, that investors care only about returns and are otherwise unmoved by the marketing efforts of investment bankers. Thus, the offer terms alone affect the probability that investors accept an offer. Consider the effect of lowering the offer price by 0.25% versus increasing the spread by 0.25%. The latter increases the probability of completion only to the extent that the bank passes some of the additional compensation on to investors. The bank's optimal strategy could involve keeping some of that compensation, even though the probability of completion would be higher if all of it were passed on to investors. In that case, underpricing makes for a more efficient vehicle to effect completion probability.

The existing empirical literature on IPO pricing does not provide a sufficiently comprehensive picture to fully test our theory. In particular, the relation between offer withdrawal and comparable firm returns leads to a truncation bias in estimates of the relation between IPO pricing and comparable firm returns.[4] Thus, we reexamine the empirical evidence in light of the predictions of our model. Consistent with prior studies, we find strong evidence that IPO prices are revised only partially in response to market returns during the waiting period. (Lowry and

[3] Loughran and Ritter (2002) also predict partial adjustment to changes in market valuations and, thus, offer similar insights regarding these time-series characteristics of IPO pricing.

[4] Busaba et al. (2001) and Benveniste et al. (2003) also recognize the implications of sample truncations for empirical analyses of IPOs. However, these studies do not examine the effects of sample truncation in the context of regressions of IPO pricing and waiting-period market returns.

Schwert, 2003, argue that, despite its statistical significance, the economic importance of partial adjustment to market returns is small after controlling for other factors that effect IPO pricing.) However, in contrast to prior studies, we find a symmetric relation between IPO price revisions and waiting-period market returns. See, for example, Loughran and Ritter, 2002; Demers and Lewellen, 2002; Lowry and Schwert, 2003). The asymmetry in the relation between IPO pricing and market returns shown in prior studies is the result of the truncated regression bias that arises from the relation between withdrawal and waiting-period market returns.

While our model predicts partial adjustment of offer prices to waiting period market returns, the economic forces modeled may not fully account for the observed tendency for offer prices to partially adjust to public information. The key consideration in addressing this question is how much issuers would be willing to pay to increase the probability of completion when comparable firm value rises. That, in turn, depends on the urgency to go public so as to exploit positive NPV projects that would otherwise be capital constrained, or to benefit from the diversification potential that a secondary market provides. Because these motives for going public are unobservable, our model can claim to be consistent only with the observed phenomenon. It does not exclude other explanations. Our model should be viewed as complementary to that of alternative theories of partial adjustment such as Loughran and Ritter (2002).

The remainder of this paper proceeds as follows. Section 2 presents a model of the role of waiting-period market returns in IPO pricing. Section 3 discusses our sample selection and data sources. Section 4 presents empirical evidence on waiting-period market returns, IPO pricing, and withdrawal. Section 5 concludes the study.

2. A Model of IPO pricing

Our model has three economic entities: the issuer, the investors, and the investment bank. Because the focus is on the role of public information in the pricing of an IPO, the model abstracts from the book-building process and presumes that the bank shops the deal to investors in a series of solicitations, each made at a specific price. This abstraction is supported by the empirical evidence in Cornelli and Goldreich (2003) and Jenkinson and Jones (forthcoming), who show that the typical indication of interest is a simple yes or no, which presumes an indicated price. If that indicated price fails, the investment bank must return to the investor because the interest shown in the first solicitation is not sufficiently precise to assess whether the market would clear at the new price. Likewise, we abstract from the advisory role of the investment bank and presume that the issuer and bank have the same model of comparable firm value. We use the term "offer" to refer to the entire sequence of solicitation rounds. These could in principle number in the dozens, but the infrequency of offer price revisions in amended filings suggests that the typical IPO involves one or perhaps a few solicitations.

Similar to Baron and Holmstrom (1980), the model presumes that investors expend time and effort in considering an IPO and that the bank's cost of shopping a deal varies inversely with its reputation for imposing costs on its investor clientele. Let $\Theta[P]$ denote the probability that making the next solicitation to investors at price P fails to attract sufficient demand. Let λ denote the bank's willingness to tolerate a failed solicitation. λ characterizes the bank's reputation according to the following three assumptions:

$$\text{The bank maintains a constant } \lambda \text{ across all offers and across each solicitation round within an offer.} \quad \text{(I)}$$

$$\text{The bank follows a policy of not permitting a solicitation price above } R_{\text{bank}}^{\max} \text{ on each round, where } R_{\text{bank}}^{\max} \text{ denotes the solution to } \Theta[R_{\text{bank}}^{\max}] = \lambda. \quad \text{(II)}$$

All agents know λ, given its permanence. $R_{\text{bank}}^{\max}$ differs across solicitation rounds.

The bank's cost function is

$$C'[\lambda] > 0 \quad \text{and} \quad C''[\lambda] > 0. \quad \text{(III)}$$

A higher λ indicates a greater expected number of solicitations required to complete the offer. This lowers investors' willingness to evaluate an offer because the expected cost of interacting with the bank is higher. Moreover, a finer partitioning of offer price revisions reduces the average difference between the final offer price and investors' reservation price, reducing investors' expected surplus. A nickel-and-diming strategy (i.e., high λ) does not necessarily mean that the bank goes back to the same investor over and over again. For example, the bank might identify subsets of investors for price experimentation. For such a strategy to work, each subset would have to be representative, which would imply that each investor would find itself in an experimental subset fairly often. Again, this implies elevated costs of dealing with that bank. Moreover, one expects that investors would tend to greet solicitations from such a bank with the response: "I'll give it a serious look after you've talked to everyone else. For now, don't waste my time." Thus, the bank's cost of finding willing investors increases with λ. As λ approaches one, investors expect endless solicitations from the bank with no possibility for surplus, implying an infinite cost to intermediate the IPO. This is captured with a convex cost function.

The bank's λ also affects issuers' demand for the bank's intermediation services. Let $N[\lambda]$ denote the volume of IPO business that the bank garners from issuers. Consider the extreme case $\lambda = 0$, i.e., the bank only makes a solicitation if there is zero probability of failure. In this case the issuer's expected surplus is at a minimum. Thus, no issuer would engage the services of the bank, i.e., $N[0] = 0$. As λ increases, so does the issuer's expected surplus, and thus, more issuers seek the bank's services, i.e., $N'[\lambda] > 0$. We further assume that $N''[\lambda] < 0$, consistent with a finite supply of issuers. Thus, the volume of offerings that the bank conducts satisfies

$$N[0] = 0, \quad N'[\lambda] > 0, \quad \text{and} \quad N''[\lambda] < 0. \quad \text{(IV)}$$

In practice, fees are stated as a percentage of offer proceeds (Chen and Ritter, 2000; Hansen, 2001). However, to simplify the analysis, we assume the following[5]

The bank's fees in the event of a completed deal are constant, F.

The bank's fees in the event of a failed deal are zero. (V)

The bank's reputation constraint is stated in Lemma 1, which outlines the driving force behind withdrawal.

Lemma 1. *The bank imposes a maximum acceptable failure probability* $\lambda < 1$ *on the issuer under Assumptions* (I)–(V).

The proof is given in the appendix. Lemma 1 establishes that the bank imposes discrete pricing limits on the issuer. Consistent with this result, empirical research finds that discreteness is a prominent characteristic of IPO pricing (Bradley et al., 2003). Given that discreteness, there comes a point in which the bank agrees to another solicitation after repeated failed solicitations only if the price is below the issuer's reservation price, i.e., $R^{\max}_{\text{bank}} < R^{\min}_{\text{issuer}}$. Thus, the bank forces withdrawal. This is stated as

Lemma 2. *There exists a finite M such that the bank's maximal acceptable solicitation price after M rejected solicitations* (*the last occurring at price* P_M), *denoted* $R^{\max}_{\text{bank}}[P_M]$, *is less than the issuer's reservation price*, $R^{\min}_{\text{issuer}}$. *As a result*, *the IPO either succeeds or is withdrawn no later than the Mth round.*

The proof is in the appendix. $R^{\min}_{\text{issuer}}$ is the value of the firm as a private entity. One of the primary reasons that firms seek an IPO is to relieve capital constraints; i.e., there are projects that the firm would take with additional capital but cannot take as a private entity. To the extent that the value of these projects is related to the value of comparable firms, the private value of the firm changes less than the public value of the firm in response to a change in comparable firm value. Our maintained hypothesis is that the issuer becomes increasingly concerned with completing the offer as comparable firm value rises, because forced withdrawal following failed solicitation(s) becomes increasingly costly. By revising their preferred offer price, denoted P^{issuer}_m, less than one-for-one with changes in comparable firm value, denoted K, the issuer increases the attractiveness of the offer to investors (i.e., $K/P_m{}^{\text{issuer}}$ increases), which increases the probability of success.

By contrast, if $\partial R^{\min}_{\text{issuer}}/\partial K = 1$, the issuer's surplus from going public is independent of comparable firm value. In that case the issuer's desire for a successful offer is unaffected by changes in comparable firm value. The issuer therefore wants to maintain the same success probability irrespective of changes in comparable-firm value, which is accomplished by adjusting P^{issuer}_m one-for-one with K. This intuition is formalized in

[5]Incorporating this feedback loop into the analysis greatly complicates matters, without contributing to the intuition of the model.

Proposition 1. *If the issuer's surplus from going public is positively related to changes in comparable firm value,*

$$\frac{\partial R_{\text{issuer}}^{\min}}{\partial K} < 1, \tag{1}$$

then the issuer's preferred offer price changes only partially in response to changes in comparable firm value,

$$\frac{\partial P_m^{\text{issuer}}}{\partial K} < 1. \tag{2}$$

The proof is in the appendix. Proposition 1 does not describe the price at which a solicitation is made to investors, only the preferences of the issuer. The bank could choose to protect its reputation by constraining the issuer to some lower price. Nevertheless, as shown in Proposition 2, this constraint does not directly lead to partial revision of offer prices in response to changes in comparable firm value: The bank's constraint moves one for one with comparable firm value. Changes in comparable firm value trigger a partial revision in the solicitation price only when the issuer's preferred price is not constrained by the bank.

Proposition 2. *If the issuer's surplus from going public is positively related to changes in comparable-firm value,*

$$\frac{\partial R_{\text{issuer}}^{\min}}{\partial K} < 1, \tag{3}$$

then the actual offer price, P_m, satisfies

$$\frac{\partial P_m}{\partial K} < 1 \text{ if and only if } P_i^{\text{issuer}} < R_{\text{bank}}^{\max}[P_{i-1}] \tag{4}$$

for some i satisfying $m \leqslant i \leqslant M$. That is, the offer price in period m is revised only partially in response to changes in K if and only if the bank does not constrain the issuer.

The proof is in the appendix. Why might a rational issue prefer to solicit the offer at a price lower than the maximum price the bank would tolerate? Such a strategy can be optimal when the issuer knows that failure at the current solicitation price causes the bank to impose a much lower price on subsequent solicitations, perhaps below the issuer's reservation price. Thus, the current solicitation could be effectively the last opportunity for the issuer to go public. Not wanting to chance it, a rational issuer prices conservatively. While a more myopic strategy of insisting on soliciting investors at the highest possible price generates higher surplus conditional on success, it yields lower surplus in expectation.

Even in those cases in which the issuer effectively sets the solicitation price (i.e., the bank's constraint that round is nonbinding), the possibility that the bank would constrain subsequent solicitation prices influences the price that the issuer chooses. Thus, the bank always exerts an important influence on pricing in our model. In particular, if the current solicitation fails, the bank could insist that subsequent solicitation occur at a price below the issuer's reservation price, thus forcing withdrawal. The issuer therefore recognizes that the current solicitation could be its last chance. This can lead to cautious pricing, particularly when the opportunity cost of a failed last-chance solicitation is high. Because comparable firm returns during the offer process proxy for that opportunity cost, they predict the degree to which the offer is underpriced.

3. Sample selection, data sources, and variable definitions

This section describes the sample of IPO firms, data sources, and proxy variables used to test our model's predictions regarding the adjustment of IPO prices to public information during the waiting period.

3.1. Sample selection

Data for IPOs are from the Thompson Financial (Securities Data Corporation) Platinum database. The initial sample includes all firms that either completed or withdrew an IPO between January 1, 1985 and December 31, 2000 (approximately 7,500 firms). We exclude from the sample banks; closed-end funds; investment trusts; master limited partnerships; American Depository Receipts; non-U.S. domiciled firms; units; offers of non voting shares; equity carve-outs and spin-offs; reverse leveraged buy outs; those filing under a small-business waiver (Securities and Exchange Commission form SB-1); those with a midpoint filing-range price of less than $5.00; and those whose shares were to be listed on exchanges other than the NYSE, AMEX, or Nasdaq. We also exclude those (39) completed offerings whose waiting periods were greater than one year, as the relevant interval for measuring comparable-firm returns for these offerings is ambiguous. Indeed, the average waiting period of over 600 calendar days for these observations suggests that some may be two separate filings recorded as one. The final sample includes 4,605 IPO filings of which 3,863 were completed and 742 were withdrawn.

Table 1 reports descriptive statistics for the sample by completion or withdrawal. The completed offerings are comparable to those of other IPO studies in terms of firm size, filing amount, Carter and Manaster (1990) underwriter ranking, venture capital backing, offer price revision, and first-day return. Of particular relevance is the waiting period, the filing date through the offer (or withdrawal) date. Completion dates are obtained from the Thompson Financial database and verified with the Center for Research in Security Prices (CRSP) database. (Completion dates are considered valid when the first trading day, according to CRSP, is within a few days of the stated completion date.) Withdrawal dates are identified in one of two ways. For 90% of the withdrawn offerings a formal withdrawal is filed (SEC form RW) and reported on the Thompson Financial database. For those 10% that do not have a formal withdrawal filing, we take the withdrawal date to be 270 days from the date of the last amendment, which is when the offering is typically deemed withdrawn by the SEC (Lerner, 1994). From Table 1, the average (median) waiting period is 70 (59)

Table 1
Descriptive statistics of initial public offerings (IPOs) from January 1985–December 2000

Data are from the Thompson Financial Securities Data Corporation Platinum database. The initial sample includes all firms that completed or withdrew an IPO between January 1, 1985 and December 31, 2000. We exclude from the sample IPOs for banks, closed-end funds, investment trusts, master limited partnerships, American Depository Receipts, non-US domiciled firms, units, non-voting shares, spin-offs, reverse leveraged buyouts, those filing under a small-business waiver (Securities and Exchange Commission form SB-1), those with a midpoint filing-range price of less than $5.00, and those whose shares were to be listed on exchanges other than NYSE, AMEX, or Nasdaq. We also exclude all (39) completed offerings whose waiting periods were greater than one year. The final sample includes 4,605 IPO filings (3,863 completed and 742 withdrawn offerings). Total equity before is the number of shares outstanding prior to the filing times the midpoint of the filing range. Filing amount is the number of shares being issued times the midpoint of the filing range. Carter and Manaster rank is a measure of underwriter quality and ranges from 1 (lowest) to 9 (highest). Waiting period is the number of calendar days between the initial filing date and the offer or withdrawal date. Offer price revision is the offer price minus the midpoint of the initial filing range, divided by the midpoint of the initial filing range. First-day return is the closing price on the day of the offering minus the offer price, divided by the offer price. Dollar and share figures are in millions except prices.

	Completed offerings		Withdrawn offerings	
Offer characteristic	Mean	Median	Mean	Median
Total equity before	$165.9	$66.5	$146.7	$70.8
Filing amount	$47.2	$29.6	$51.5	$33.0
Shares filed	3.46	2.50	3.81	2.70
Midpoint filing price	$11.76	$12.00	$12.49	$12.00
Carter and Manaster rank	7.3	8.1	7.2	8.0
Waiting period (days)	70	59	269	216
Issue proceeds	$47.5	$28.8		
Shares sold	3.33	2.50		
Offer price	$11.99	$12.00		
Offer-price revision	1.9%	0.0%		
First-day return	23.2%	8.7%		

calendar days for completed offerings and 269 (216) for withdrawn offerings. During a period of this duration a potential exists for substantial variation in issuing-firm value.

3.2. *Proxies for information flow during the waiting period*

The literature has identified three types of information flow during the waiting period: public, private, and spillover. To focus on public information, our empirical analysis controls for the effects of both private and spillover information on IPO pricing.

3.2.1. *Public information: The returns of comparable firms*

The market valuation of publicly traded comparable firms is readily observable in real time. While the flow of public information about the IPO firm's value includes

Table 2
Comparable-firm Returns During the Waiting Period

The mean and standard deviation of comparable firm returns during the waiting period. The waiting period encompasses the initial filing date through the offer or withdrawal date. Comparable firms are identified for each sample firm from the Center for Research in Security Prices (CRSP) universe using the 48 industries defined in Fama and French (1997) modified as follows: (1) software firms (Standard Industry Classification (SIC) codes 7370-7375) are removed from computers and business services and assigned to a new industry "software", (2) bank holding companies (SIC code 6799) are removed from financials, (3) pharmaceuticals (SIC 8731) are removed from business services and assigned to pharmaceuticals, (4) Internet firms (identified from Securities Data Corporation business descriptions) are assigned to a new industry, Internet, with comparable firm returns obtained using an equally weighted average of the Goldman Sachs and Street.com internet indices. We compute the cumulative equal-weighted return for each sample firm's comparable-firm portfolio during the sample firm's waiting period. Statistics are reported for 27 industries having at least 1% of the sample.

FF industry	Number of filings	Percent of sample	Waiting-period return	
			Mean	Standard deviation
Apparel	56	1.2%	0.0%	8.9%
Building maintenance	53	1.2	5.7	10.9
Business services	272	6.0	4.9	14.9
Chips	326	7.2	4.8	17.9
Computers	150	3.3	3.8	19.4
Construction materials	59	1.3	3.5	14.4
Electrical equipment	58	1.3	4.8	10.2
Energy	65	1.4	5.1	14.4
Entertainment	107	2.4	0.6	11.2
Financial	66	1.5	4.0	8.0
Healthcare	199	4.4	3.5	15.2
Household	57	1.3	6.4	13.4
Insurance	105	2.3	5.0	8.2
Internet	253	5.6	9.7	34.4
Lab equipment	88	1.9	6.5	20.8
Machine	85	1.9	3.9	9.3
Meals	94	2.1	2.2	9.4
Medical equipment	199	4.4	3.3	14.8
Personal services	61	1.3	3.0	13.1
Pharmaceutical	334	7.4	5.7	21.6
Recreational products	51	1.1	1.9	9.5
Retail	281	6.2	1.3	8.1
Software	613	13.5	5.8	23.3
Telecommunication	215	4.7	7.2	22.1
Transportation	100	2.2	2.8	10.6
Utilities	59	1.3	4.7	7.2
Wholesale	170	3.7	2.7	11.3
Full sample	4,605	100	4.5	17.9
CRSP equal-weighted index			4.0	10.1

additional information such as news reports or industry trade data, we limit our proxy for public information flow to the market valuation of comparable firms. We form comparable firm portfolios by assigning firms from the CRSP universe to the

48 industry classes as defined by Fama and French (1997), with modifications noted in Table 2. We then compute the cumulative equal-weighted return with monthly rebalancing for each IPO firm's comparable firm portfolio over the IPO firm's waiting period.

From Table 2, considerable variation is found in comparable firm value during the waiting period. For example, the standard deviation of comparable firm returns is 18%. Empirical studies of IPO pricing typically use a market index to proxy valuation changes, which captures roughly half of the variation in valuation implied by industry comparable firms (for example, the standard deviation of CRSP equal-weighted index returns during the waiting period is 10%.) Furthermore, the average comparable firm return during the waiting period is 4.5%, whereas the average offer price revision is only 1.9% (Table 1).

3.2.2. Spillover information: The pricing of other offerings

Benveniste et al. (2003) argue that information from the pricing of other IPOs cannot be regarded as exogenous to the intermediation process: Private information that intermediaries obtain about the demand for one IPO is likely informative about the demand for other IPOs. Thus, we distinguish between the spillover information from the pricing of other offerings and purely public information from the returns of comparable firms.

We construct two spillover measures. The first is the average offer price revision of IPOs completed during the 30 calendar days prior to offer or withdrawal date. The second is the average first-day return for those IPOs. The focus is on recent offerings, not all IPOs completed during the waiting period, to capture the state of the primary market at the time of the offering or withdrawal. To isolate spillover information, which is endogenous to the investment banking process, from that which is clearly public information, we orthogonalize these measures with respect to the waiting-period comparable firm return, issue size, the Carter and Manaster underwriter ranking, and venture capital backing.

3.2.3. Private information: Offer-price revisions

A third source of information during the waiting period is the book building process, whereby underwriters acquire private information about demands from investors. Book building theories argue that information endogenous to the intermediation process requires compensation in the form of a partial revision in the offer price. Previous studies use the revision in offer price from filing to offer date to proxy for such information. However, as pointed out in Loughran and Ritter (2002), empirical evidence shows that offer prices are revised in response to public as well as private information. No compelling economic argument can be made that investors should be compensated when public information triggers a change in the offer price. Thus, as with the spillover variables, we orthogonalize this measure with respect to the public information variables.

4. Empirical evidence on IPO pricing and public information

This section replicates the analysis of IPO pricing and waiting period market returns of prior studies using a sample of completed offerings. We then confirm that IPO withdrawal is also related to waiting period market returns, and thus, analyses of IPO pricing restricted to completed offerings may be subject to truncation bias. Finally, we present new evidence on the relation between IPO pricing and waiting period market returns using the full sample of completed and withdrawn offerings.

4.1. Inferences from a sample of completed offerings

Table 3 replicates the analysis of the relation between IPO pricing and information flow during the waiting period found in the literature. As in much of the literature, the specification is ordinary least squares (OLS) restricted to a sample of completed offerings. In Columns 1 and 2 the dependent variable is the revision in offer price. In Columns 3 and 4 the dependent variable is the first-day return.

Studies dating back to Logue (1973) suggest that offer prices are revised only partially in response to waiting-period market returns. Column 1 of Table 3 confirms this finding using comparable firm portfolio returns. The coefficient for comparable firm returns is 0.41 (standard error of 0.04). By contrast, the average post-IPO beta against comparable firms (not presented in Table 3) is 1.2 (standard error of 0.36). This partial adjustment of IPO prices to public information is the central prediction of our model.

The specification in Column 2 allows the slope of the relation between offer price revisions and comparable firm returns to differ with respect to positive versus negative comparable firm returns. Consistent with several studies, the adjustment of IPO prices to public information appears to be more complete when comparable firm returns are negative (coefficient $0.34 + 0.33 = 0.67$) versus positive (coefficient $= 0.34$). This inference of larger response on the downside stems from a truncated regression bias.

Partial adjustment of offer prices to waiting-period market returns implies predictable first-day returns. Columns 3 and 4 of Table 3 report coefficient estimates from regressions of first-day returns on comparable firm returns during the waiting period. The results of Table 3 are consistent with prior studies. From Column 3, the estimated coefficient for comparable firm returns is 0.51 (standard error of 0.07). In column 4 we allow the slope of the relation to differ with respect to positive versus negative comparable firm returns. The estimated slope is significantly attenuated when comparable firm returns are negative versus positive, as indicated by the coefficient on the negative-returns interactive variable, $R^{\text{comps}-}$. In particular, the estimated coefficient on comparable-firm returns is 0.58 when those returns are positive, versus 0.16 (0.58–0.42) when negative.

4.2. IPO withdrawal and sample selection bias

The relation between IPO pricing and information flow during the waiting period can be estimated only with completed offerings, as offer prices are not observed for

Table 3
Offer-price revision and first-day return regressions (ordinary least squares)

Coefficient estimates, with standard errors in parenthesis, from ordinary least squares (OLS) regressions of offer price revisions (Columns 1 and 2) and first-day returns (Columns 3 and 4) on comparable firm returns during the waiting period. Offer price revision is the offer price minus the midpoint of the initial filing range, divided by the midpoint of the initial filing range. First-day return is the closing price on the day of the offering minus the offer price, divided by the offer price. R^{comps} is the return of a comparable firm portfolio (see Table 2) during the waiting period. R^{comps-} is equal to R^{comps} when negative and 0 otherwise. Carter and Manaster rank is a measure of underwriter quality and ranges from 1 (lowest) to 9 (highest). Filing amount is the log of the ratio of the midpoint of the initial filing range times the number of shares filed divided by the total market value of all Center for Research in Security Prices (CRSP) firms at the time of filing, divided by 1,000,000. Venture backed is equal to one if the firm received venture capital prior to the filing date and 0 otherwise. Spillover revisions and Spillover returns are the average offer-price revision and average first-day return of all initial public offerings (IPOs) completed during the 30 days prior to the offering, orthogonalized with respect to the other independent variables. Bookbuild is the orthogonalized offer price revision for the sample firm. Bookbuild^{-} is equal to Bookbuild when negative and zero otherwise.

Independent variables	Dependent variable			
	Offer price revisions		First-day returns	
	(1)	(2)	(3)	(4)
R^{comps}	0.408 (0.038)	0.342 (0.043)	0.507 (0.071)	0.575 (0.099)
R^{comps-}		0.332 (0.097)		−0.417 (0.189)
Filing amount	−0.038 (0.006)	−0.040 (0.006)	−0.109 (0.009)	−0.102 (0.009)
Venture backed	0.022 (0.007)	0.025 (0.007)	0.075 (0.011)	0.058 (0.011)
Carter and Manaster rank	0.022 (0.007)	0.023 (0.007)	0.041 (0.003)	0.039 (0.003)
Spillover revisions	0.473 (0.054)	0.462 (0.053)	0.606 (0.095)	0.615 (0.095)
Spillover returns	0.222 (0.036)	0.246 (0.036)	0.871 (0.083)	0.765 (0.086)
Bookbuild			1.036 (0.065)	1.360 (0.128)
Bookbuild^{-}				−0.786 (0.177)
Adjusted R^2	20.5	20.8	46.1	47.5

withdrawn offers. From Table 1, 16% of IPO filings are withdrawn. According to our model, both IPO pricing and withdrawal are related to comparable firm returns during the waiting period, implying that a sample of completed offerings is nonrandom with respect to the regressors in the pricing equations. This leads to a truncated distribution of regression disturbances and therefore biased coefficient estimates. (A similar comment could be made at the filing stage: Filing for an IPO is a nonrandom event, so the sample of IPO filings is truncated to just those firms that have something to gain from an IPO.)

Table 4 provides evidence on the relation between offer withdrawal and comparable firm returns. A substantial difference exists in the duration of the waiting period for completed versus withdrawn offerings (see Table 1). This induces a spurious positive relation between waiting-period returns and withdrawal. Returns tend to drift positive and the sample period 1985–2000 was one of the strongest periods in U.S. market history. Thus, a longer waiting period tends to be associated with larger comparable firm returns. Moreover, the decision to withdraw likely occurs long before the formal filing of withdrawal, if such a filing is even made. This makes precise dating of the relevant interval for measuring comparable firm returns difficult, particularly in those cases in which the inferred waiting period is long. We use the SEC 270-day rule as an implicit cutoff for withdrawal and assign those observations with waiting periods greater than 270 days the median waiting period of those that filed within 270 days, 123 days.

Busaba et al. (2001) and Benveniste et al. (2003) show a negative relation between the frequency of offer withdrawal and the market return during the waiting period. In Panel A of Table 4 we show a similar result using comparable firm returns. Thus, consistent with our model, both the pricing and the withdrawal of IPOs is related to comparable firm returns during the waiting period. In addition, our evidence shows that withdrawal is concentrated in the left tail of the distribution of comparable firm returns. From Panel A of Table 4, the frequency of withdrawal is relatively high, 32%, when comparable firm returns are substantially below zero (lowest return quartile) and relatively low, about 12%, otherwise (highest three return quartiles). Thus, the truncation bias is only sensitive to comparable firm returns when those returns are negative.[6]

To outline the effect of sample truncation on regression coefficient estimates, Eq. (5) presents a statement of the process generating an offer price revision, ΔP^{offer}

$$\Delta P^{\text{offer}} = a^{\text{revision}} + b^{\text{revision}} R^{\text{comps}} + \varepsilon^{\text{revision}}, \tag{5}$$

where $\varepsilon^{\text{revision}}$ is a mean zero disturbance that is, by construction, independent of R^{comps}. Eqs. (6) and (7) present a statement of the process generating offer withdrawal

$$I_{\text{complete}} = \begin{cases} 1 & \text{if } z^* > 0 \\ 0 & \text{if } z^* \leqslant 0 \end{cases} \tag{6}$$

[6]The evidence in Table 4 suggests that the appropriate location for a kink in the comparable firm returns coefficient is at the first quartile of returns, not zero. However, we locate the kink at zero to be consistent with the literature.

Table 4
Withdrawal and comparable firm returns

Panel A reports withdrawal frequency by comparable firm return quartile. Sample firms are assigned to quartiles based on comparable firm returns during the waiting period. The waiting period of offers withdrawn more than 270 days is set equal to the median (123 days from initial filing date) of firms that file a withdrawal statement within 270 days. Panel B reports probit analysis of initial public offering (IPO) withdrawal. The dependent variable is equal to one for withdrawn offerings and zero for completed offerings. Maximum likelihood coefficient estimates are reported with standard errors in parenthesis. R^{comps} is the comparable-firm return during the waiting period. R^{comps} equals R^{comps} when negative and zero otherwise. Carter-Manaster Rank is a measure of underwriter quality and ranges from 1 (lowest) to 9 (highest). Filing amount is the log of the ratio of the midpoint of the initial filing range times the number of shares filed divided by the total market value of all Center for Research in Security Prices (CRSP) firms at the time of filing, divided by 1,000,000. Spillover revisions and Spillover returns are the average offer price revision and first-day return of all initial public offerings (IPOs) completed during the 30 days prior to the offering, respectively, orthogonalized with respect to the other independent variables. Spillover withdrawals is the number of offers withdrawn within 123 days of filing relative to all other active IPOs during the 30 days preceding the offer or withdrawal date.

Quartile	Comparable-firm return	Percent withdrawn
Panel A		
1	−12.3%	31.6%
2	−0.7	11.7
3	4.8	11.1
4	21.2	12.0

Independent variables:	Dependent variable: withdrawal
Panel B	
R^{comps}	0.463
	(0.192)
$R^{\text{comps}-}$	−5.154
	(0.436)
Filing amount	0.279
	(0.036)
Carter and Manaster rank	−0.085
	(0.014)
Spillover revisions	−1.000
	(0.296)
Spillover returns	0.241
	(0.152)
Spillover withdrawals	70.193
	(13.014)

and

$$z^* = a^{\text{complete}} + b^{\text{complete}} R^{\text{comps}} + \varepsilon^{\text{complete}}, \tag{7}$$

where I_{complete} is an indicator function for a completed deal, z^* is a latent variable for completion probability, and $\varepsilon^{\text{complete}}$ is a mean zero disturbance that is, by construction, independent of R^{comps}.[7] According to the theory developed in Section 2, offers are priced so as to cause b^{complete} to be positive.

Consider an OLS estimate of Eq. (5). Denote the estimated pricing coefficient $\hat{b}^{\text{revision}}$:

$$E\left[\hat{b}^{\text{revision}}\right] = \frac{\text{cov}[R^{\text{comps}}, \Delta P^{\text{offer}}]}{\text{var}[R^{\text{comps}}]} = b^{\text{revision}} + \frac{\text{cov}[R^{\text{comps}}, \varepsilon^{\text{revision}}]}{\text{var}[R^{\text{comps}}]}. \tag{8}$$

The estimate $\hat{b}^{\text{revision}}$ is unbiased in a random sample, because the disturbance $\varepsilon^{\text{revision}}$ is independent of R^{comps}. However, in a sample selected on the basis of offer completion $\hat{b}^{\text{revision}}$ is likely biased because $\varepsilon^{\text{revision}}$ is not random with respect to offer completion. This can be seen in two steps.

First, suppose that two offerings, A and B, experience the same R^{comps} over their waiting periods and are similar in all other respects, except that offer A is priced relatively aggressively for idiosyncratic reasons, i.e., $\varepsilon^{\text{revision}}(A) > \varepsilon^{\text{revision}}(B)$. Our model predicts that aggressive pricing increases the probability that the investment bank forces withdrawal, i.e.,

$$\text{Cov}(\varepsilon^{\text{revision}}, \varepsilon^{\text{complete}}) < 0. \tag{9}$$

Second, the conditional mean of $\varepsilon^{\text{complete}}$ in a sample of completed offers exceeds the unconditional mean (i.e., zero) because firms with low $\varepsilon^{\text{complete}}$ tend to not make it into the sample. Because a lower R^{comps} raises the minimum $\varepsilon^{\text{complete}}$ consistent with completion, this truncation effect is greatest with low or negative R^{comps}. At higher values of R^{comps}, most of the $\varepsilon^{\text{complete}}$ distribution is consistent with completion and there is little truncation. That is,

$$\text{cov}\left[R^{\text{comps}}, \varepsilon^{\text{complete}} | I_{\text{complete}} = 1\right] < 0, \tag{10}$$

For sufficiently high R^{comps}, truncation becomes irrelevant, so

$$\text{cov}\left[R^{\text{comps}}, \varepsilon^{\text{complete}} | I_{\text{complete}} = 1\right] \cong 0 \tag{11}$$

for high values of R^{comps}.

Eqs. (9) and (10) imply that R^{comps} positively covaries with $\varepsilon^{\text{revision}}$ in a sample of completed offerings, inducing a positive bias in the OLS coefficient estimate $\hat{b}^{\text{revision}}$. Further, the bias is relatively severe at low values of R^{comps} and relatively immaterial at high values of R^{comps}. Thus, OLS estimates of $\hat{b}^{\text{revision}}$ exhibit a kink, even if there is no asymmetry in the true relation between offer price revisions and R^{comps}. By similar logic, OLS estimates of the relation between first-day returns and R^{comps} are negatively biased when restricted to a sample of completed offerings.

Let $f(R^{\text{comps}})$ represents the relation between the conditional expectation of $\varepsilon^{\text{revision}}$ given a completed offer and R^{comps} as described above, where $f'(R^{\text{comps}}) > 0$. That is, in a sample of completed offerings $\varepsilon^{\text{revision}}$ follows

$$\varepsilon^{\text{revision}} = f(R^{\text{comps}}) + \eta, \tag{12}$$

[7]The relations are not necessarily linear, with zero intercept, as specified. That is for analytical convenience.

where η is orthogonal to R^{comps}. Heckman's two-step selection correction procedure involves adding a proxy for $f(R^{\text{comps}})$ into the regression, denoted λ, which is derived from a probit regression of withdrawal on R^{comps}. Similar to adding a correlated omitted variable to a misspecified regression, adding λ causes the true coefficient on R^{comps} to be estimated with less bias, or no bias if λ completely captures the in-sample covariance between $\varepsilon^{\text{revision}}$ and R^{comps}.

4.3. Inferences from a sample of completed and withdrawn offerings

This section reports estimates of the relation between IPO pricing and public information using Heckman's two-step selection correction procedure. In the first-step probit regression, the dependent variable is zero if the offer is completed and one if the offer is withdrawn. The independent variables include those used in the pricing regressions of Table 3, with the addition of an identifying variable, the withdrawal rate of other offerings during the 30 days prior to each offer or withdrawal date. The withdrawal rate is equal to the number of offerings withdrawn within 123 days of their initial filing date divided by the number of active registrants. We restrict the measure to firms that withdrew within 123 days of their initial filing date to capture the current IPO market climate.[8] Table 4, panel B reports maximum likelihood estimates from the probit analysis. The coefficient on R^{comps} is relatively close to zero whereas the coefficient on $R^{\text{comps}-}$ is -5.15 (standard error of 0.44). Thus, as in Panel A, a negative relation is shown between withdrawal and comparable firm returns that is driven primarily by the higher incidence of withdrawals when comparable firms are performing poorly.

Table 5 reports estimates from the second step of the Heckman procedure: an OLS regression that parallels the regressions of Table 3, except that the instrument for withdrawal probability from the probit analysis, λ, is added as an independent variable. The effect of selection bias on inferences regarding the relation between IPO pricing and public information can be seen by comparing the estimates of Table 3 and Table 5.

From Column 2 of Table 5, the coefficient estimate for R^{comps} in the offer price revision regression, 0.37 (standard error of 0.05), is similar to the corresponding estimate from Table 3, 0.34 (standard error of 0.04). However, the difference in the coefficient estimates for $R^{\text{comps}-}$ is striking. From Column 2 of Table 5, the coefficient estimate is 0.03 (standard error of 0.20), whereas the coefficient estimate from Table 3 is 0.33 (standard error of 0.10). Thus, no statistically distinguishable asymmetry exists in the pricing regression when λ is included as a regressor, implying that the asymmetry in the OLS coefficient estimate for R^{comps} is the result of sample truncation bias. We find similar evidence of sample-truncation bias in the first-day return regressions. From Column 4 of Table 5, the coefficient estimate on R^{comps}, 0.54 (standard error of 0.10), is similar to the corresponding estimate from Table 3, 0.58 (standard error of 0.10), while the coefficient estimate on $R^{\text{comps}-}$ from Table 5

[8]One hundred twenty-three days is the median number of days between initial filing and withdrawal for those offers withdrawn within 270 days, the point at which the offer is deemed withdrawn by the SEC.

Table 5
Offer price revision and first-day return regressions (Heckman two-step selection correction estimation)

Coefficient estimates, with standard errors in parenthesis, are from Heckman two-step selection correction regressions of offer-price revisions (columns 1 and 2) and first-day returns (Columns 3 and 4) on comparable-firm returns during the waiting period. Offer-price revision is the offer price minus the midpoint of the initial filing range, divided by the midpoint of the initial filing range. First-day return is the closing price on the day of the offering minus the offer price, divided by the offer price. R^{comps} is the return of a comparable firm portfolio (see Table 2) during the waiting period. $R^{\text{comps}-}$ is equal to R^{comps} when negative and 0 otherwise. Carter and Manaster rank is a measure of underwriter quality and ranges from 1 (lowest) to 9 (highest). Filing amount is the log of the ratio of the midpoint of the initial filing range times the number of shares filed divided by the total market value of all Center for Research in Security Prices (CRSP) firms at the time of filing, divided by 1,000,000. Venture backed is equal to one if the firm received venture capital prior to the filing date and 0 otherwise. Spillover revisions and Spillover returns are the average offer-price revision and average first-day return of all initial public offerings (IPOs) completed during the 30 days prior to the offering, orthogonalized with respect to the other independent variables. Bookbuild is the orthogonalized offer price revision for the sample firm. Bookbuild$^-$ is equal to Bookbuild when negative and zero otherwise. λ is the probability of withdrawal estimated in the first step using the probit analysis of Table 4.

Independent variables	Dependent variable			
	Offer-price revision		First-day return	
	(1)	(2)	(3)	(4)
R^{comps}	0.367	0.365	0.587	0.543
	(0.034)	(0.046)	(0.088)	(0.099)
$R^{\text{comps}-}$		0.029		0.063
		(0.196)		(0.325)
Filing amount	−0.027	−0.028	−0.128	−0.124
	(0.006)	(0.006)	(0.010)	(0.013)
Venture backed	0.026	0.026	0.069	0.057
	(0.007)	(0.007)	(0.011)	(0.011)
Carter and Manaster rank	0.019	0.019	0.048	0.045
	(0.008)	(0.008)	(0.004)	(0.004)
Spillover revisions	0.386	0.391	0.738	0.737
	(0.056)	(0.058)	(0.095)	(0.113)
Spillover returns	0.268	0.266	0.795	0.719
	(0.036)	(0.036)	(0.079)	(0.080)
Bookbuild			1.044	1.356
			(0.049)	(0.125)
Bookbuild$^-$				−0.781
				(0.170)
λ	0.127	0.118	−0.215	−0.206
	(0.032)	(0.064)	(0.070)	(0.101)
Adjusted R^2	20.9	20.8	46.3	47.6

of 0.06 (standard error of 0.33) differs substantially from the corresponding coefficient estimate from Table 3 of −0.42 (standard error of 0.19). Again, the difference between the coefficient estimates of Tables 5 and 3 can be attributed to the significant negative coefficient on λ.

The coefficient estimates for the other regressors in Table 5 and Table 3 are largely unaffected by the inclusion of λ. This is consistent with the relatively weak association between withdrawal probability and these other regressors. From Panel A of Table 4, the difference in withdrawal frequency for firms in the first quartile of comparable firm returns versus firms in the fourth quartile of comparable firm returns is 20%. Though not tabulated, the difference in withdrawal frequency for firms sorted on the Carter and Manaster (1990) underwriter ranking and offer size is 8% and −5%, respectively. We cannot assess the potential bias in the coefficient estimate for ΔP^{offer} because that regressor cannot be included in the probit regression.[9]

Table 5 is useful in assessing the relative importance of public versus private information in explaining IPO pricing. Suppose that two pieces of information arrive during the IPO waiting period: a public signal in the form of a 10% increase in the valuation of comparable firms, which causes a 3.7% increase in the offer price (see Table 5, Column 1); and a private signal (orthogonal to comparable firm returns) that causes a 3.7% increase in the offer price. Thus both signals have the same effect on the offer price. According to Table 5, Column 3, the public signal generates a 5.9% increase in the IPO's first-day return and the private signal generates a 3.8% (3.7%∗1.044) increase in the IPO's first-day return. Thus, the adjustment of offer prices to public signals is evidently less complete than the adjustment to private signals of similar magnitude. This conclusion regarding the importance of partial adjustment to public information contrasts with Lowry and Schwert (2002). The key differences in our methodologies is that we orthogonalize the proxy for private information (the offer price revision) with respect to public information and use returns of industry matched comparable firms as opposed to a broad market index to proxy for public information. When the private information regressor is not orthogonalized and the public information proxy is a broad market index, much of the signal in comparable firm returns appears via the offer price revision regressor.

The evidence in Loughran and Ritter (2003) suggests a structural shift in the pricing of IPOs in late 1998. For example, the average first-day return quadrupled from the 1990–1998 period to the 1999–2000 period. This shift gives rise to two concerns. First, our results might be driven by the unusual post-1998 sub period. Second, differences across regimes might be confounding our cross-sectional estimates. In results not tabulated we repeated the analysis in Tables 3–5 separately for the pre-1998 (1985–1998) and post-1998 (1999–2000) periods. We find evidence of nonstationarities in the coefficients of the control variables, as many of the

[9]For similar reason, one cannot determine whether the statistically significant kink on ΔP^{offer} in the Table 5, Column 4 regression is the result of a sample selection bias or true economic effects. However, the parellel between ΔP^{offer} and R^{comps} (both represent information about IPO surplus arriving during the waiting period) suggests that selection bias is the culprit.

coefficients differ significantly across these two samples. However, the coefficients for comparable firm returns for the pre-1998 sample are qualitatively the same as that for the full sample, while the coefficients for comparable firm returns for the post-1998 sample are too noisy to draw reliable inferences from. Thus, our inferences regarding IPO pricing and comparable firm returns do not stem from the unusual post-1998 sub period, and they are not driven by the structural shift in underpricing.

5. Conclusion

The partial adjustment of IPO prices to information learned during the waiting period has been the subject of considerable research, both empirical and theoretical. Most of this research has focused on partial adjustment to private information. The partial adjustment to public information has received comparatively little attention, despite the two effects having roughly similar magnitude.

Our explanation for partial adjustment to public information is based on simple intuition: Issuers facing an uncertain outcome want the offer priced to maximize the expected surplus from going public. This involves a consideration of both offer proceeds and the probability of completion. The pricing implications drawn from our model explain several empirical regularities associated with IPOs in addition to the partial-adjustment phenomenon. In particular, the model predicts the well-documented unconditional underpricing of IPOs. Also because underpricing is related to a common factor and waiting periods overlap, the model can explain hot issues markets. The model also predicts the inverse relation between withdrawal frequency and comparable firm returns. This last prediction has important implications for the proper specification of the empirical relation between IPO pricing and comparable firm returns. In particular, we show that the asymmetry in the relation between IPO pricing and comparable firm returns explained in prior studies is an econometric illusion caused by a truncated regression bias. Consistent with our model, a properly specified test indicates a symmetric partial adjustment of IPO prices to comparable firm returns during the waiting period.

While the model does not directly address partial adjustment to private information, the model's intuition is suggestive. Any information that causes the issuer to perceive greater surplus associated with going public should lead to less aggressive pricing (i.e., more severe underpricing). If the book-building process yields private information about the firm's potential surplus from a public offering, the rational issuer wants a partial adjustment of the offer price to that information so that the probability of success is tilted in the desired direction. Thus, the model's predictions are complementary to existing book-building models, which focus on the need for partial adjustment to compensate investors for divulging private information.

Our model is also consistent with the common complaint of issuers, that the bank was coercive in handling the offering. Suppose that the bank fails to generate sufficient interest in soliciting an offering at a price that the issuer had agreed was acceptable. Dragging investors through yet another round of solicitation could cause them to conclude that working with that bank (or that banker) is a waste of their time and effort. Thus, the bank could insist upon a low offer price to ensure a high probability of success and thus preserve its reputation capital. The issuer walks away feeling as if it were coerced into giving away the firm, when, in fact, further solicitation at a high price would simply be a negative net present value proposition for the bank.

Finally, we make no claims that our model fully explains partial adjustment of offer prices to public information. Our model should be viewed as complementary to that of alternative theories of partial adjustment such as Loughran and Ritter (2002). We emphasize instead that our model is largely consistent with several aspects of observed IPO pricing beyond partial adjustment to public information.

Appendix A

Proof of Lemma 1. The bank maximizes profit

$$\max\ N[\lambda] \cdot (F \cdot (1-\lambda) - C[\lambda]), \tag{A.1}$$

which implies the first order condition

$$\frac{F \cdot (1-\lambda) - C[\lambda]}{F + C'[\lambda]} = \frac{N[\lambda]}{N'[\lambda]}. \tag{A.2}$$

Given assumption (IV), the right-hand side denominator in Eq. (A.2) decreases with λ and the numerator increases. Thus the right-hand side starts at zero and increases in λ thereafter. The left-hand side denominator increases with λ and the left-hand side numerator decreases. Thus, the left-hand side is decreasing in λ. At $\lambda = 1$ (100% chance of failure) costs are maximal and expected fee revenue is nil, so the left-hand side is negative. If F does not exceed $C[0]$, then the left-hand side is always negative because costs are minimal and fee revenue is certain (equal to F) when $\lambda = 0$. In that case, the first-order condition is never satisfied and the bank does not agree to undertake the offering. Therefore, we consider the case in which the left-hand side takes on positive values ($F > C$) for at least some λ. In that case, the first-order condition is satisfied at some $0 < \lambda < 1$. Differentiating again and using $F \cdot (1-\lambda) - C[\lambda] > 0$ (first-order condition) yields the second-order condition for a maximum.

Proof of Lemma 2. Write the conditional distribution of investors' reservation price as $\Theta[P_m; P_{m-1}, K]$, where all solicitations prior to round m failed with the last occurring at the price P_{m-1}, and K denotes the comparable-firm valuation of the IPO firm. Let ρ_{issuer} denote the inverse Cumulative Distribution Function of $R^{\min}_{\text{issuer}}$ using $\Theta[P; \infty, K]$, i.e.,

$$\Theta\left[R^{\min}_{\text{issuer}}; \infty, K\right] = \rho_{\text{issuer}}; \quad 0 < \rho_{\text{issuer}} < 1. \tag{A.3}$$

The bank makes an initial solicitation to investors at price P if $\Theta[P;\infty,K]<\lambda$. Let the function $P^{\text{bank}}[P_{\text{prior}}]$ denote the value of P^{bank} that solves

$$\Theta\left[P^{\text{bank}};P_{\text{prior}},K\right]=\lambda. \tag{A.4}$$

The probability of failure in the first round at the constrained price is $P^{\text{bank}}[\infty]=\lambda$. The conditional probability of failure in the mth round, given failure in rounds 1 through $m-1$, is

$$\frac{\Theta[P_m;\infty,K]}{\Theta[P_{m-1};\infty,K]}=\lambda<1. \tag{A.5}$$

Set $M=\text{int}(\log(\rho_{\text{issuer}})/\log(\lambda))+1$. If $m>M$, then

$$\Theta[P_m;\infty,K]<\Theta[P_M;\infty,K]<\lambda^{\log(\rho_{\text{issuer}})/\log(\lambda)}=(e^{\log(\lambda)})^{\log(\rho_{\text{issuer}})/\log(\lambda)}=\rho_{\text{issuer}}, \tag{A.6}$$

which implies $P_m<R_{\text{issuer}}$. Hence, the issuer rejects the bank's mth round constraint.

Proof of Proposition 1. We make one additional assumption regarding Θ:

$$\Theta[P_1;\infty,K]=\Theta[z] \quad \text{where } z=P_1-K, \tag{VI}$$

so that investors' demand for the IPO remains unchanged when the offer price is revised one-for-one with changes in comparable firm value. If the offer price is revised less than one-for-one, investor demand increases following positive changes in comparable firm value and decreases following negative changes in comparable firm value. The proof of Proposition 1 involves a backward recursive argument starting in the terminal round, whose existence is established in Lemma 2. Denote

$$\pi_m^*=\max_{P_m}\ \Theta[P_m;P_{m-1},K]F_m+(1-\Theta[P_m;P_{m-1},K])P_m, \tag{A.7}$$

where F_m is the expected IPO proceeds given that solicitation m just failed:

$$F_m=\begin{cases}\pi_{m+1}^* & \text{if } m<M,\\ R_{\text{issuer}}^{\min} & \text{if } m=M.\end{cases} \tag{A.8}$$

π_m^* is the issuer's expected IPO proceeds just prior to the mth solicitation, assuming that optimal choices are made. Two additional lemmas establish Proposition 1:

Lemma A.1.

$$\frac{\partial P_m^{\text{issuer}}}{\partial K}<1 \quad \text{if and only if} \quad \frac{\partial F_m}{\partial K}<1. \tag{A.9}$$

and

$$\frac{\partial P_m^{\text{issuer}}}{\partial K} = 1 \quad \text{if } \frac{\partial F_m}{\partial K} = 1. \tag{A.10}$$

Lemma A.2.

$$\text{If } \frac{\partial R_{\text{issuer}}^{\min}}{\partial K} < 1 \text{ then } \frac{\partial F_m}{\partial K} < 1 \quad \forall\, m \leqslant M. \tag{A.11}$$

Proof of Lemma A.1. The first-order condition is defined by differentiating Eq. (A.7):

$$\Theta \frac{\partial F_m}{\partial P_m} + (P_m^{\text{issuer}} - F_m) \cdot \left(-\frac{\partial \Theta}{\partial P_m}\right) + (1 - \Theta) = 0. \tag{A.12}$$

Note that

$$\frac{\partial F_M}{\partial P_M} = \frac{\partial R_{\text{issuer}}}{\partial P_M} = 0 \quad \text{and} \quad \frac{\partial F_m}{\partial P_m} = 0 \quad \forall\, m < M. \tag{A.13}$$

The first follows because the issuer's reservation price depends on the private valuation, not on its IPO bidding behavior. The second follows because $F_m = \pi_{m+1}^*$ is an optimized value on P_m. Thus, the first-order condition for π_m^* , Eq. (A.12), is

$$(P_m^{\text{issuer}} - F_m) \cdot \left(-\frac{\partial \Theta}{\partial P_m}\right) + (1 - \Theta) = 0. \tag{A.14}$$

The second order condition for a maximum is

$$(P_m^{\text{issuer}} - F_m) \cdot \frac{\partial^2 \Theta}{\partial P_m^2} + 2 \frac{\partial \Theta}{\partial P_m} > 0. \tag{A.15}$$

Totally differentiating the first order condition yields

$$\frac{\mathrm{d}P_m^{\text{issuer},U}}{\mathrm{d}K} = \frac{(P_m^{\text{issuer},U} - F_m)(\partial^2\Theta/\partial P_m^2) + (1 + (\partial F_m/\partial K)) \cdot (\partial\Theta/\partial P_m)}{(P_m^{\text{issuer},U} - F_m)(\partial^2\Theta/\partial P_m^2) + 2(\partial\Theta/\partial P_m)}. \tag{A.16}$$

Thus,

$$\frac{\mathrm{d}P_m^{\text{issuer}}}{\mathrm{d}K} < 1 \quad \text{if and only if } \frac{\partial F_m}{\partial K} < 1 \quad \text{and} \quad \frac{\mathrm{d}P_m^{\text{issuer}}}{\mathrm{d}K} = 1 \quad \text{if } \frac{\partial F_m}{\partial K} = 1. \tag{A.17}$$

Proof of Lemma A.2. The proof is by mathematical induction. Suppose that $\partial F_{m+1}/\partial K < 1$, and show that it implies $\partial F_m/\partial K < 1$. $\partial F_M/\partial K = \partial R_{\text{issuer}}/\partial K < 1$ holds as a condition of the proposition.

- If the $m+1$ solicitation is made at P^{bank}_{m+1}, then

$$F_m = \pi_{m+1} = \lambda F_{m+1} + (1-\lambda)P^{\text{bank}}_{m+1}. \tag{A.18}$$

Differentiate the identity $\mathrm{d}/\mathrm{d}K\,\Theta[P^{\text{bank}}; P_{\text{prior}}, K] = 0$ yields

$$\frac{\mathrm{d}P^{\text{bank}}}{\mathrm{d}K} = -(\partial\Theta/\partial K)/(\partial\Theta/\partial P^{\text{bank}}) = 1. \tag{A.19}$$

Thus, the bank's constraint moves one-for-one with comparable firm value. Using this,

$$\frac{\partial F_m}{\partial K} = \lambda\,\frac{\mathrm{d}F_{m+1}}{\mathrm{d}K} + (1-\lambda)\,\frac{\mathrm{d}P^{\text{bank}}_{m+1}}{\mathrm{d}K} = 1 - \lambda\left(1 - \frac{\mathrm{d}F_{m+1}}{\mathrm{d}K}\right), \tag{A.20}$$

so, $\mathrm{d}F_{m+1}/\mathrm{d}K < 1$ implies $\mathrm{d}F_m/\mathrm{d}K < 1$.

- If the $m+1$ solicitation is made at P^{issuer}_{m+1}, then

$$F_m = \pi^*_{m+1} = \Theta F_{m+1} + (1-\Theta)P^{\text{issuer}}_{m+1} \tag{A.21}$$

and

$$\frac{\partial F_m}{\partial K} = \Theta\,\frac{\mathrm{d}F_{m+1}}{\mathrm{d}K} - (P^{\text{issuer}}_{m+1} - F_{m+1})\left(\frac{\partial\Theta}{\partial K} + \frac{\partial\Theta}{\partial P_{m+1}}\frac{\mathrm{d}P^{\text{issuer}}_{m+1}}{\mathrm{d}K}\right) + (1-\Theta)\frac{\mathrm{d}P^{\text{issuer}}_{m+1}}{\mathrm{d}K}. \tag{A.22}$$

From the induction assumption and Lemma 3, $\mathrm{d}P^{\text{issuer}}_{m+1}/\mathrm{d}K < 1$, so the bracketed term is negative.

Using Eq. (A.20), Eq. (A.22) becomes

$$\frac{\partial F_m}{\partial K} = \Theta\,\frac{\mathrm{d}F_{m+1}}{\mathrm{d}K} + (1-\Theta)\cdot 1, \tag{A.23}$$

which defines a convex combination of one and something less than one ($\mathrm{d}F_{m+1}/\mathrm{d}K$). Hence, the partial derivative is less than one.

Proof of Proposition 2. $R^{\max}_{\text{bank},m}$ solves the identity $\Theta[R^{\max}_{\text{bank},m}; P_{m-1}, K] = \lambda$. Equivalently,

$$\frac{\Theta[R^{\max}_{\text{bank},m}; \infty, K]}{\Theta[P_{m-1}; \infty, K]} = \lambda. \tag{A.24}$$

Differentiating this identity with respect to K,

$$\frac{\partial\Theta[R^{\max}_{\text{bank},m}; \infty, K]}{\partial K}\cdot\left(1 - \frac{\mathrm{d}R^{\max}_{\text{bank},m}}{\mathrm{d}K}\right) = \lambda\,\frac{\partial\Theta[P_{m-1}; \infty, K]}{\partial K}\cdot\left(1 - \frac{\mathrm{d}P_{m-1}}{\mathrm{d}K}\right) \tag{A.25}$$

To maintain the identity Eq. (A.24), irrespective of K, it must be that

$$\frac{\partial\Theta[R^{\max}_{\text{bank},m}; \infty, K]}{\partial K} = \lambda\,\frac{\partial\Theta[P_{m-1}; \infty, K]}{\partial K} \tag{A.26}$$

or

$$\left(1-\frac{\mathrm{d}R_{\mathrm{bank},m}^{\max}}{\mathrm{d}K}\right)=\left(1-\frac{\mathrm{d}P_{m-1}}{\mathrm{d}K}\right)\Rightarrow\frac{\mathrm{d}R_{\mathrm{bank},m}^{\max}}{\mathrm{d}K}=\frac{\mathrm{d}P_{m-1}}{\mathrm{d}K}. \tag{A.27}$$

Now we turn to the proposition, beginning with the "only if" part. If the bank constrains in round $m-1$ (i.e., the prior period solicitation price $P_{m-1}=R_{\mathrm{bank},m-1}^{\max}$), then

$$\frac{\mathrm{d}R_{\mathrm{bank},m}^{\max}}{\mathrm{d}K}=\frac{\mathrm{d}P_{m-1}}{\mathrm{d}K}=\frac{\mathrm{d}R_{\mathrm{bank},m-1}^{\max}}{\mathrm{d}K}=\frac{\mathrm{d}P_{m-2}}{\mathrm{d}K}. \tag{A.28}$$

The first and third equalities are from Eq. (A.27). If the bank constrains in all earlier rounds,

$$\frac{\mathrm{d}R_{\mathrm{bank},m}^{\max}}{\mathrm{d}K}=\frac{\mathrm{d}R_{\mathrm{bank},1}^{\max}}{\mathrm{d}K}=1. \tag{A.29}$$

The second equality comes from totally differentiating the unconditional distribution of Θ. Thus, if the bank constrains, then the offer price moves one-for-one with K.

Suppose instead that the bank does not constrain in one or more rounds preceding round m. Let $i<m$ denote the latest round in which the constraint is not binding. If the constraint is not binding then, $\mathrm{d}P_i/\mathrm{d}K<1$ from Lemma 3 and Eq. (3) applied to Lemma A.2. Applying Eq. (A.27) and the fact that round $i+1$ is constrained, by construction of i, yields

$$\frac{\mathrm{d}P_i}{\mathrm{d}K}<1\Rightarrow\frac{\mathrm{d}R_{\mathrm{bank},i+1}^{\max}}{\mathrm{d}K}<1. \tag{A.30}$$

Also, because round $i+1$ is constrained, $P_{i+1}=R_{\mathrm{bank},i+1}^{\max}$. Thus $\mathrm{d}P_{i+1}/\mathrm{d}K<1$. Continuing to apply Eq. (A.27) gives $\mathrm{d}R_{\mathrm{bank},i+2}^{\max}/\mathrm{d}K<1$ and, eventually, $\mathrm{d}R_{\mathrm{bank},m}^{\max}/\mathrm{d}K<1$.

The "if" part follows. Consider proof by contradiction. Assume that $\mathrm{d}P_m/\mathrm{d}K<1$ but that all prior rounds are constrained. The second assumption is inconsistent with the first by the preceding analysis. Thus, if $\mathrm{d}P_m/\mathrm{d}K<1$, it must be that the constraint is not binding in some earlier period.

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