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How Do Diversity of Opinion and Information Asymmetry Affect Acquirer Returns?

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We examine the theoretical predictions that link acquirer returns to diversity of opinion and information asymmetry. Theory suggests that acquirer abnormal returns should be negatively related to information asymmetry and diversity-of-opinion proxies for equity offers but not cash offers. We find that this is the case and that, more strikingly, there is no difference in abnormal returns between cash offers for public firms, equity offers for public firms, and equity offers for private firms after controlling for one of these proxies, idiosyncratic volatility. (*JEL* G31, G32, G34)

This article examines whether variables suggested by diversity-of-opinion models and information asymmetry models are helpful in understanding the cross-sectional variation in acquirer announcement returns using a sample of pure equity offers and pure cash offers for public and private firms from 1980 to 2002. We document that these variables, the uncertainty proxies, explain a significant fraction of the cross-sectional variation in acquirer announcement returns. Perhaps most strikingly, after controlling for the uncertainty proxies, there is no difference in abnormal returns between cash offers for public firms, equity offers for public firms, and equity offers for private firms.

Using two proxies for diversity of opinion employed previously in the literature, the standard deviation of analyst forecasts and breadth of ownership, we show that bidder abnormal returns for acquisitions of

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public firms paid for with equity are lower the higher the diversity of opinion. The economic significance of the relation between diversity of opinion and abnormal returns is substantial. For instance, going from a bidder with a low standard deviation of analyst forecasts (one standard deviation below the mean) to a bidder with a high standard deviation of analyst forecasts (one standard deviation above the mean) reduces the announcement abnormal return by roughly 2.6%. In contrast, there is no negative relation between bidder abnormal returns and diversity of opinion for acquisitions of private firms paid for with equity or for acquisitions of public firms paid for with cash.

A firm's idiosyncratic volatility can proxy for information asymmetry. We find this variable to be extremely helpful in understanding acquirer abnormal returns as predicted by information asymmetry models. In regressions explaining acquirer returns for acquisitions of public firms paid for with equity, the abnormal return falls as idiosyncratic volatility increases. When the proxies for diversity of opinion are added to regressions that already include idiosyncratic volatility as an explanatory variable, they are not significant. Finally, acquirer abnormal returns for acquisitions of public firms paid for with cash increase as bidder idiosyncratic volatility increases.

Though our results are supportive of the role of proxies for diversity of opinion and information asymmetry as determinants of bidder abnormal returns, we also find results that are difficult to reconcile with diversity-of-opinion models. We find at best limited support for the prediction of these models that larger acquisitions paid for with equity should have a worse impact on bidders with greater diversity of opinion. Further, diversity-of-opinion models cannot explain why we find some evidence that bidder returns increase with diversity of opinion for cash offers for public firms. In contrast, the predictions of information asymmetry models hold across offer types.

The article proceeds as follows. In Section 1, we review the theories that motivate our uncertainty proxies and summarize the predictions of these theories. We describe our sample of acquisitions and acquiring firms in Section 2. In Sections 3, 4, and 5, we examine how variables that proxy for diversity of opinion, information asymmetry, and resolution of uncertainty help explain acquirer abnormal returns. In Section 6, we explore whether these variables can explain the differences in abnormal returns across types of acquisitions. We investigate further the robustness of our results in Section 7. We conclude in Section 8.

1. Hypotheses

In this article, we investigate the relation between bidder returns and proxies for diversity of opinion and information asymmetry (the uncertainty proxies). In this section, we briefly review and draw implications about acquirer returns from the theoretical models in which the uncertainty proxies play a key role. Before we do so, it is useful to point out that in a rational expectations model with normally distributed returns, the absolute expected return conditional on the sign of the return increases with the volatility of the return. Hence, all else equal, in such a model the expected return conditional on bad news would be more negative for a stock that is more volatile and the expected return conditional on good news would be more positive for such a stock.[1] This simple mechanism could explain the existence of a relation between abnormal returns and the uncertainty proxies. However, the models we rely on for our empirical work predict both the sign of the announcement return and its relation with the uncertainty proxies.

1.1 Diversity of opinion

Miller (1977), Chen, Hong, and Stein (2002), and Hong, Scheinkman, and Xiong (2006), among others, develop models in which diversity of opinion about a firm's prospects leads to a downward-sloping demand curve for its stock. With these models, the slope of the demand curve increases with diversity of opinion among investors. As the supply of shares available for trading (float) increases, it has to be absorbed by investors who have a lower opinion of the stock. These models therefore imply that acquisition announcements by firms with greater diversity of opinion should have worse returns when the acquisition increases the bidder's float.

A more direct prediction of diversity-of-opinion models is that the adverse impact of an increase in the float increases with diversity of opinion. If, for an equity offer for a public firm, the size of the offer represents the size of the increase in the bidder's float, we expect bidder returns to decrease in the proxy for diversity of opinion interacted with the proportional increase in the float. However, not all newly issued shares necessarily add to the float. In particular, as emphasized by Baker, Coval, and Stein (2006), some target shareholders may be sleepy, so that the shares they receive do not really add to the float. In this case, for a given increase in the supply of shares, we would expect the bidder abnormal return to fall with bidder diversity of opinion and in the proportion of target shareholders who are not sleepy.

Cash acquisitions have no impact on the float. The impact on float of acquisitions of private firms paid for with equity depends on if and when the owners of the acquired firm sell the shares. If they sell all their shares immediately, the impact on float is similar to acquisitions of public firms paid for with equity. This seems unlikely, because the owners may be

[1] Diamond and Verrecchia (1987) explicitly analyze returns conditional on the arrival of positive and negative news.

prevented from doing so with lock-up agreements, they may want to be influential in the acquiring firm, and they may have capital gains that make it suboptimal for them to sell the shares. We provide evidence consistent with the hypothesis that the float increases less with acquisitions of private firms paid for with equity than with acquisitions of comparable public firms.

We use two proxies for diversity of opinion: dispersion of analyst forecasts and breadth of ownership. Recent literature uses the dispersion of analyst forecasts as a measure of diversity of opinion, while Chen, Hong, and Stein (2002) propose a model in which diversity of opinion is negatively related to breadth of ownership.[2,3]

1.2 Information asymmetry

A traditional explanation for the negative bidder announcement returns for acquisitions of public firms paid for with equity, put forward by Travlos (1987) and inspired by Myers and Majluf (1984), is that the announcement signals to the market that bidder management believes the firm's common stock is overvalued. We therefore expect bidder abnormal returns to be negative for equity offers. When management makes a cash offer, the market infers that equity is worth more than its market value, which is good news and leads to higher abnormal returns. With equity acquisitions of private firms, the seller can obtain confidential information directly from the acquirer, so the acquirer would not expect to benefit by using overpriced equity as a means of payment for such acquisitions. It could even be that the willingness of the seller to receive equity is better news for acquirers with greater information asymmetries since the seller can help certify that the acquiring firm is not overvalued.[4] Consequently, we expect either no relation or a positive relation between abnormal returns and the proxies for information asymmetry for acquisitions of private firms paid for with equity.

Krasker (1986) extends the Myers and Majluf (1984) model to show that there is a negative relation between the post-issue price and the size of the equity issue when management can choose the size of the issue. Krasker's model therefore suggests a negative relation between abnormal returns and the size of an acquisition.

An additional prediction of the information asymmetry models is that, everything else equal, the expected growth should be lower if the firm pays for the acquisition of a public firm with equity as opposed to cash.

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[2] See, for instance, Diether et al. (2002); Diether (2004), and Scherbina (2003).

[3] We are grateful to the referee for suggesting the use of this measure.

[4] This argument is made in the context of private equity placements by Hertzel and Smith (1993). It would not apply if the seller expects to sell his shares immediately. The management of a public target can also acquire private information. Consequently, the argument is valid only if, in that case, management cannot provide a certification benefit, perhaps because of conflicts of interest.

Dierkens (1991) explores the relation between abnormal returns for equity issues and proxies for the nature of the information environment. Her proxies are the standard deviation of the earnings announcement abnormal return, the firm's idiosyncratic volatility, the firm's turnover, and the number of public announcements of the firm. We use the first two of these as proxies for information asymmetry in our tests. One of these proxies, the idiosyncratic volatility of the stock, has also been used in the literature recently as a measure of diversity of opinion.[5]

1.3 Resolution of uncertainty

Typically, one would expect uncertainty to get resolved more for firms with a higher level of uncertainty about growth prospects. We therefore have to make sure that our uncertainty proxies are not just proxies for resolution of uncertainty since existing models suggest that resolution of uncertainty could be associated with worse abnormal returns for acquirers. In the models of McCardle and Viswanathan (1994) and Jovanovic and Braguinsky (2004), an acquisition signals adverse information about the bidder's prospects and resolves uncertainty about these prospects. Pástor and Veronesi (2006) and Johnson (2004) show that uncertainty about a firm's long-term growth prospects increases firm value in an efficient market. This effect is stronger for firms with better growth prospects. On the basis of these models, we expect an acquisition announcement that reduces uncertainty about a firm's long-term growth prospects to be associated with a drop in firm value unless there is an accompanying synergy gain large enough to offset the resolution-of-uncertainty effect. This prediction of the resolution-of-uncertainty models should hold for all types of offers.

1.4 Summary of model predictions

Table 1 summarizes the predictions of the models when applied to acquirer returns. We include the predictions associated with the resolution-of-uncertainty hypothesis even though our focus is on diversity of opinion and information asymmetry models. As we previously argued, the resolution-of-uncertainty hypothesis is important because we need to ensure that our uncertainty proxies are not significant because they are correlated with uncertainty resolution. The models predict that, for equity offers for public firms, acquirer abnormal returns are negatively related to the uncertainty proxies. However, each of the models has unique predictions for cash offers of public firms and equity offers of private firms. We can use these predictions to see which model, if any, better explains bidder returns. In particular, the diversity-of-opinion models predict a negative relation between abnormal returns for equity offers for public firms but not

[5] See Boehme, Danielsen, and Sorescu (2006) for references.

Table 1
Model predictions

Increase in	Acquirer abnormal returns: Acquisition of public firms paid for with equity	Acquisition of public firms paid for with cash	Acquisition of private firms paid for with equity
Diversity of opinion	Decrease	No effect	No effect
Information asymmetry	Decrease	Increase	Increase or no effect
Resolution of uncertainty	Decrease	Decrease	Decrease

for other acquisitions. The information asymmetry models make opposite predictions for cash and equity offers for public firms. The resolution-of-uncertainty models make the same prediction for all types of acquisitions. In addition, the diversity-of-opinion models make strong predictions for the role of the size of the offer and the composition of target shareholders. The information asymmetry models have implications for changes in expected growth associated with the type of financing for acquisitions of public firms. We investigate these various additional predictions.

2. The Data

We first describe the sample of acquisitions and then turn to the characteristics of bidders and targets in our sample. Finally, we introduce our proxies for diversity of opinion, information asymmetry, and uncertainty resolution.

2.1 The sample of acquisitions

To analyze the relation between the uncertainty proxies and the acquirer's acquisition announcement abnormal return, we start from a sample of successful and unsuccessful acquisition announcements constructed from the Securities Data Company's (SDC) US Mergers and Acquisitions Database. Our sample is restricted to pure cash and pure equity offers to avoid complications that arise when considering mixed offers. Since there are too few pure cash offers for private firms, our sample is limited to pure cash and pure equity offers for public firms and pure equity offers for private firms. None of the models we consider has predictions that would make it helpful to consider mixed offers. We require that the deal value corresponds to at least 1% of the market value of the assets of the acquirer (defined as the book value of assets minus the book value of equity plus the market value of equity). In addition, the sample of acquisitions meets the following criteria:

1. The acquisition attempt is announced in the period from 1980 to 2002 and neither the acquirer nor the target has another merger announcement in the three-day window;

2. The acquirer controls less than 50% of the shares of the target at the announcement date and a successful acquirer obtains 100% of the target shares;
3. The deal value is equal to or greater than $1 million;
4. The target is a US public firm or a US private firm;
5. Data on the acquirer is available from CRSP and COMPUSTAT;
6. The deal is classified by SDC as either successful, unconditional, or withdrawn;
7. If successful, the deal is completed in less than 1000 days.

We find 4322 acquisition announcements that meet these criteria. Next, for tests that require analyst data, an acquiring firm must have a forecast for long-term growth of earnings per share the month preceding the acquisition (month−1) and be followed by at least three analysts at that time so that it is meaningful to compute a standard deviation of long-term forecasts.[6] Our information on analyst forecasts is obtained from the Summary History File of the Institutional Brokers Estimate System (I/B/E/S) database. This requirement leaves a subset with analyst data of 1553 announcements.

2.2 Bidder characteristics

Table 2 provides information on acquirer and deal characteristics for the full sample of acquisitions and the subset with analyst data. Our regressions use the same control variables as Moeller, Schlingemann, and Stulz (2004). These variables are in italics in the table. Panel A of Table 2 shows that imposing the requirement of analyst forecasts availability increases sharply the mean and median transaction value for the acquisitions considered. Further, the acquisitions become less important relative to the market value of equity of the acquirer or the market value of the assets. It follows from this comparison that the sample of acquisitions with analyst forecasts is not a representative sample of all acquisitions: La Porta (1996); Hong, Lim, and Stein (2000), and Diether, Malloy, and Scherbina (2002), among others, note that the intersection of CRSP, COMPUSTAT, and I/B/E/S is severely skewed towards larger companies. Panel B shows that, whether using book value of assets, market value of assets, or market value of equity, acquirers with analyst forecasts are much larger than acquirers without such forecasts. The acquirers with forecasts also have lower leverage. We use the market-to-book ratio, computed as total assets minus the book value of common equity plus the market value of common equity divided by total assets, as a proxy for Tobin's q. There is evidence

[6] To compute a standard deviation, we need at least two observations, but we focus on cases where we have at least three forecasts to avoid putting too much weight on outliers. Though we believe that this measure is more reliable, our results also hold if we require only two analyst forecasts to compute our measure of dispersion.

that the acquirers with forecasts are valued more than those without, as can be seen from the higher Tobin's q for acquirers with forecasts. There is no difference between the two groups of firms for operating cash flow and prior stock performance.

The next two variables reported in Panel B are proxies for information asymmetry. The idiosyncratic volatility measure, labeled volatility in the tables, is the standard deviation of the residuals from a market model regression estimated from 205 days before the announcement to six days before the announcement. The earnings residual standard deviation is measured as the standard deviation of all three-day cumulative abnormal returns around earnings announcements from I/B/E/S using the market

Table 2
Sample summary statistics

	All (n = 4322)		With analyst data (n = 1553)	
	Mean	Median	Mean	Median
Panel A: Deal characteristics				
Transaction value (TV)	591.50	47.09	1347.05[a]	150.00[a]
TV/Market value of equity (MVE)	0.6104	0.1811	0.2961[a]	0.1147[a]
Days to completion	110	96	100[a]	85[a]
Competed ×100%	4.19		3.86	
Cash in consideration ×100%	15.97		14.36	
Equity in consideration ×100%	84.05		85.64	
Hostile ×100%	2.48		2.19	
Tender offer ×100%	7.17		8.24	
Same industry ×100%	32.92		35.48[c]	
Public target ×100%	53.91		61.30[a]	
Private target ×100%	46.09		38.70[a]	
(Public target | equity) ×100%	37.95		46.94[a]	
(Private target | equity) ×100%	46.09		38.70[a]	
(Public target | cash) ×100%	15.96		14.36	
Panel B: Acquirer characteristics				
Cash/book value of assets	0.2040	0.1090	0.2121	0.1272[c]
Book value of assets	3403.382	280.847	7421.24[a]	770.217[a]
Market value of equity	2501.7	295.836	5705.417[a]	1198.72[a]
Leverage	0.3741	0.2832	0.3040[a]	0.2063[a]
Tobin's q	2.8427	1.6734	3.6046[a]	2.1638[a]
Operating cash flow	0.1514	0.0942	0.1480	0.1166[b]
Small dummy ×100%	44.52		11.59[a]	
Run-up ×100%	14.54	−4.55	18.41	−0.70[a]
Governance index ×100%	90.49		82.36[a]	
Period 1998–2000 × 100%	29.06		36.06[a]	
Liquidity index ×100%	17.33	7.98	21.09	10.61[a]
Volatility	0.0346	0.0288	0.0297[a]	0.0262[a]
Earnings residual (std)	0.0678	0.0556	0.0692	0.0575[b]
Breadth of ownership	0.0188	0.0082	0.0339[a]	0.0204[a]
Panel C: Abnormal returns				
$CAR_{(-1,+1)}$ × 100%—all	0.8184	−0.3477	−0.6437[a]	−0.8450[a]
$CAR_{(-1,+1)}$ × 100%—public | equity	−2.2815	−2.0385	−2.834	−2.4230
$CAR_{(-1,+1)}$ × 100%—private | equity	3.4232	0.9028	1.8907[a]	0.8780
$CAR_{(-1,+1)}$ × 100%—public | cash	0.6664	−0.1048	−0.3153[b]	−0.3110
CAR (public | equity—private | equity)	−5.7047*	−2.9413*	−4.7247*	−3.3010*
CAR (public | equity—public | cash)	−2.9479*	−1.9337*	−2.5187*	−2.1120*
CAR (private | equity—public | cash)	2.7568*	1.0076*	2.2060*	1.1890*

Table 2
(Continued)

	I/B/E/S firms $n = 439{,}774$ (1)	Sample firms $n = 1553$ (2)	Firm median $n = 1506$ (3)	Differences (1)—(2)	(1)—(3)	(2)—(3)
Panel D: I/B/E/S long-term growth (LTG) forecasts						
LTG (std)	4.328	5.075	4.722	-0.747^{a}	-0.394^{b}	0.353^{a}
	[3.000]	[3.250]	[3.690]	$[-0.250]^{a}$	$[0.690]^{a}$	$[0.080]^{b}$
LTG (median)	15.875	23.330	21.919	-7.455^{a}	-6.044^{a}	1.411^{a}
	[14.000]	[20.000]	[20.000]	$[-6.000]^{a}$	$[-6.000]^{a}$	$[0.000]^{a,-}$
Revision in LTG (std)	0.027^{a}	−0.038	−0.012	0.065	0.039^{a}	−0.026
	$[0.000]^{a}$	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]
Revision in LTG (median)	-0.172^{a}	−0.017	−0.009	−0.155	-0.163^{a}	−0.008
	$[0.000]^{a}$	[0.000]	[0.000]	$[0.000]^{c,+}$	[0.000]a,+	[0.000]

The table presents a sample of successful and unsuccessful acquisitions by publicly listed US acquirers obtained from the SDC Merger and Acquisition Database for the period 1980–2002. The sample includes all deals involving US private targets with 100% equity payment and public targets with either 100% equity or 100% cash payment. The italicized variables are control variables in Moeller, Schlingemann, and Stulz (2004). The subsample with analyst data includes acquirers with long-term growth analyst forecasts by three or more analysts. In Panel A, the *transaction value* is from SDC and represents the total value of consideration paid by the acquirer, excluding fees and expenses. The *market value of equity* is for the fiscal year end prior to the announcement. The *market value of assets* is the book value of assets minus the book value of equity plus the market value of equity. *Days to completion* is the number of days between the announcement and effective date (for successful deals). *Competed, hostile, tender offer, and cash and equity in consideration* are from SDC. *Same industry* deals involve targets with the same two-digit SIC code as that of the bidder. In Panel B, *cash* includes cash and marketable securities. *Leverage* is measured as the ratio of long-term and short-term debt to the market (book) value of assets. *Tobin's q* is defined as the ratio of the market value of assets to the book value of assets. *Operating cash flow* is defined as sales minus the cost of goods sold, sales and general administration, and working capital change, normalized by the book value of assets. *Small dummy* is equal to one if the acquirer has a market value of equity equal to or less than the market value of equity of the smallest quartile of NYSE firms in the year of the acquisition. *Run-up* is measured as the market-adjusted buy-and-hold return over the period from 205 days to six days prior to the announcement of the deal. *Governance* is a dummy variable equal to one if the reported Governance Index from Gompers, Ishii, and Metrick (2003) for the acquirer is above the sample median. *Period 1998–2000* is a dummy variable equal to one if the deal is announced during the calendar years 1998–2000. The *liquidity index* for the target is calculated as the value of corporate control transactions in the two-digit SIC code for each year divided by the total book value of assets of firms in the two-digit SIC code for that year. *Volatility* is the standard deviation of the market-adjusted residuals of the daily stock returns measured during the period starting from 205 to six days prior to the acquisition announcement. *Earnings Residual* (*std*) is the standard deviation of all three-day cumulative abnormal returns around earnings announcements from I/B/E/S using the market model over the 5-year period preceding the acquisition announcement. *Breadth of ownership* of the acquirer is the fraction of mutual funds that own the stock in the quarter prior to the acquisition. Variables in italics are used as control variables in the regression analysis. In Panel C, the $\text{CAR}_{(-1,+1)}$ denotes the three-day cumulative abnormal return (in percent) measured using market model residuals. In Panel D, the standard deviation and median of the long-term growths forecasts and forecast revisions are reported in column (1) for all I/B/E/S firms with three or more analysts and available data on long-term earnings growth forecasts, in column (2) for the sample firms the month prior to the announcement, and in column (3) for the median of the time series for sample firms using all available months excluding the months prior, during, and after the announcement of the deal. The forecast revisions for the sample firms are the difference from the month before the announcement to the month after the announcement. Forecast revisions for all I/B/E/S firms and firm median are measured over a two-month period with overlapping windows. Superscripts a, b, and c denote, respectively, statistical significance at the 1, 5, and 10% levels based on t-tests (means) and Wilcoxon-tests (medians) of the unpaired differences between the two samples in Panels A through C and for Panel D across the groups in columns (1)–(2) and (1)–(3) and a paired t-test (sign-rank test) is used for the mean (median) paired difference between (2)–(3). For Panel C, *, denotes significance at the 1% level for the difference in means or medians between the abnormal returns of the subsamples containing private targets with 100% equity payment and public targets with either 100% equity or 100% cash payment. For Panel D, $^{-}$ and $^{+}$ respectively denote a negative and positive test statistic in case the mean or median paired difference is rounded to zero in the table, yet is significantly different from zero.

model over the five years preceding the acquisition announcement. Though idiosyncratic volatility is lower for firms with analyst data, the earnings residual standard deviation is not. The last variable shown in Panel B, breadth of ownership, is a proxy for diversity of opinion. As diversity of opinion about a stock increases, more mutual funds would prefer to sell the stock short if they were not prevented from doing so by short-sale restrictions. Consequently, an increase in diversity of opinion is associated with a decrease in the breadth of ownership, which we measure as the fraction of mutual funds that own the stock in the quarter prior to the acquisition. Breadth of ownership is greater for firms with analyst data.

We use the Center for Research in Securities Prices (CRSP) database to collect daily return data for our sample of acquirers and data for the equally weighted index. We estimate the acquirer abnormal returns, $CAR_{(-1,+1)}$, associated with the three-day window surrounding the acquisition announcements in our sample for each year using standard event study methods (see, e.g., Brown and Warner (1985)). We compute market model abnormal returns using the CRSP equally weighted index, where the parameters for the market model are estimated over the (−205, −6) day interval. In Panel C of Table 2, we report the mean and median announcement abnormal returns. Not surprisingly, we find that the abnormal returns differ significantly across offer types for the whole sample. These significant differences are preserved when we consider the restricted sample, but the mean announcement abnormal return of firms with analyst data is roughly 150 basis points less than the mean announcement abnormal return of firms in the unrestricted sample. The differences in acquirer size between the two samples (see Moeller, Schlingemann, and Stulz (2004)) as well as the greater prevalence of equity offers in the restricted sample help explain this difference in abnormal returns.

Another selection bias induced by restricting the sample to only acquirers with at least three analyst forecasts is that the fraction of acquirers with analyst forecasts increases over time. Consequently, the percentage of acquisitions made by firms followed by at least three analysts is higher in recent years, so the proportion of acquisitions included in our sample is higher on average during the last five years of the sample period. The restricted sample includes 21.55% of the unrestricted acquisitions from 1980 through 1990, 38.50% from 1991 through 2002, and 62.53% from 1998 through 2002. The latest merger wave is therefore overrepresented in the restricted sample.

2.3 Measures of diversity of opinion and uncertainty resolution constructed from analyst forecasts

Our empirical work using analyst forecasts focuses on the long-term earnings growth forecast, which I/B/E/S defines as a three to five year forecast of the expected annual increase in operating earnings over the company's next full business cycle. The primary reason we choose the long-term growth forecast, instead of quarterly or yearly forecasts, is that it features prominently in valuation models. This long-term forecast also has several other advantages. First, quarterly or yearly earnings forecasts are affected by how close a firm is to the end of a fiscal quarter or year and by how important earnings guidance is for a firm. These considerations are less likely to influence the long-term growth forecast. Second, quarterly or yearly forecasts typically have to be normalized to be made comparable across firms and the normalization may introduce noise in comparisons of forecasts across firms. Because the long-term forecast is an expected growth rate, it is directly comparable across firms.

The main variable of interest in our analysis is the dispersion of analyst forecasts before the acquisition announcement (month-1) measured by the standard deviation of these forecasts.[7] The difficulty with this variable is that the dispersion of analyst forecasts is somewhat higher when a firm has few—but more than one—analysts. Therefore, we also use a different measure of dispersion of analyst forecasts. For each number of analyst forecasts, we rank the standard deviation of forecasts and we call high analyst dispersion firms those that rank in the top decile of dispersion of analyst forecasts among firms with the same number of analyst forecasts.

Panel D of Table 2 provides information on our analyst measures. All of our data come from the Summary File. We show the mean and median of the standard deviation and the median of the long-term growth forecasts. The first column is the whole I/B/E/S sample for which a long-term growth forecast is available and there are at least three analysts. The second column provides data for the sample of acquisitions.

Comparing the acquirers to all I/B/E/S firms using a Wilcoxon median test, we first see that the acquirers have higher long-term growth prospects and more dispersion in long-term growth forecasts. The third column in Panel D of Table 2 is the full time-series median of the sample of acquiring firms excluding the forecasts from one month before to one month after the merger announcement. The overall median of dispersion and levels of long-term growth forecasts across all acquirers measured during the month before the acquisition are lower and the same as, respectively, for all other months. However, the mean value of the long-term growth

[7] The month of the acquisition (month 0) is defined as the I/B/E/S statistical period in which the announcement occurs unless that announcement is within six business days of the end of the period. For announcements that occur within six business days of the end of the period, the next month is considered the month of the acquisition because only the forecasts of the next month are expected to be affected by the acquisition (see Pound (1988)). There is a risk that this procedure will lead to a misclassification of some forecasts as having been made before the announcement when in fact they were made after the announcement. Though we report results using this classification method, our results hold if, instead, we use month–2 for the acquisitions announced during the last six business days of the I/B/E/S statistical period.

analyst forecasts dispersion for acquirers is higher for the month before an acquisition than for the other months. The last two rows of the table show that neither the standard deviation of long-term growth forecasts nor the median long-term growth forecasts change significantly from before to after the announcement.

3. Diversity of Opinion and Abnormal Returns

We use multiple regressions to evaluate whether diversity-of-opinion proxies are helpful in explaining acquirer abnormal returns. All regressions use industry-fixed effects at the two-digit SIC code. We first discuss the results for acquisitions of public firms paid for with equity. In Section 3.2, we investigate whether information about diversity of opinion of the target is useful to explain acquirer abnormal returns. We turn to other offers in Section 3.3.

3.1 Diversity of opinion and acquisitions of public firms paid for with equity

All the regressions in the following tables have the same format. We regress the three-day acquirer abnormal return on a constant, one or more uncertainty proxies, and control variables:

$$\begin{aligned}\text{Abnormal return} = \text{Constant} + \sum_{i=1}^{n} \beta_i \text{Uncertainty proxy}_i \\ + \sum_{j=n+1}^{N} \beta_j \text{Control variable}_j + \varepsilon \qquad (1)\end{aligned}$$

In regression (1) of Panel A of Table 3, the uncertainty proxy is the measure of analyst forecasts dispersion.[8] The regression has no control variables. We find that the measure of analyst forecasts dispersion has a significant negative coefficient. The coefficient is −0.0028. The standard deviation of the measure of dispersion of analyst forecasts is 5.64. Consequently, a difference of two standard deviations of the measure of dispersion of analyst forecasts corresponds to an abnormal return difference of 3.2%.

To account for known determinants of acquisition abnormal returns, we add in regression (2) variables that the literature often uses to explain acquirer abnormal returns. As discussed earlier, we use the same variables as Moeller, Schlingemann, and Stulz (2004). To save space, we do not report the coefficients on these variables.[9] In addition, to take into account

[8] We also estimate regressions to which we add the square of the dispersion of analyst forecasts. We find no evidence to support such a specification.

[9] Tables with these coefficients are available from the authors.

Table 3
Cross-sectional regressions

Panel A: Public targets and equity payment	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
LTG (std)	−0.0028[b]	−0.0023[c]			−0.0020		−0.0026[b]	
	0.024	0.063			0.200		0.048	
LTG (std) × TVMVE					−0.0007			
					0.788			
LTG (std) × TVMVE × Institutional ownership							0.0014	
							0.606	
Top decile LTG (std)			−0.0315[b]					
			0.021					
Breadth of ownership				0.1064[c]		0.0813		0.0142
				0.060		0.221		0.832
Breadth of ownership × TVMVE						0.105		
						0.576		
Breadth of ownership × TVMVE × Institutional Ownership								0.7550[b]
								0.019
Mutual fund holdings				−0.1263[b]		−0.0134		−0.0838
				0.031		0.865		0.169
Mutual fund holdings × TVMVE						−0.2447[c]		
						0.069		
Mutual fund holdings × TVMVE × Institutional Ownership								−0.2263[c]
								0.084
Institutional ownership							0.0011	−0.0055
							0.929	0.678
TVMVE		−0.0253[a]	−0.0253[a]	−0.0174[a]	−0.0228	0.0187	−0.0280[a]	−0.0115
		0.006	0.005	0.004	0.118	0.317	0.005	0.103
Constant	0.0551[c]	0.0456	0.0358	0.0093	0.0445	−0.0029	0.0454	0.0115
	0.088	0.168	0.290	0.371	0.184	0.798	0.182	0.284
Observations	724	720	720	1502	720	1502	720	1502
Adjusted R^2	0.025	0.070	0.070	0.046	0.068	0.051	0.068	0.054

(continued overleaf)

Table 3
(Continued)

Panel B: Private and cash deals	Private targets | Equity (1)	(2)	(3)	Public targets | Cash (4)	(5)	(6)
LTG (std)	−0.0009			0.0025*		
	0.380			0.112		
Top decile LTG (std)		0.0041*			0.0254*	
		0.735*			0.231	
Breadth of ownership			−0.1684*			−0.1226[c]*
			0.219*			0.056
Mutual fund holdings			−0.0042			0.0886*
			0.959			0.157
TVMVE	0.0156*	0.0157*	0.0100*	0.0063*	0.0063*	0.0037*
	0.255	0.253	0.202	0.575	0.590	0.160
Constant	0.0307*	0.0098*	−0.009*	−0.0423[c]	−0.0281*	0.0315
	0.204	0.666	0.706	0.085	0.276	*0.261*
Observations	599	599	1689	219	219	604
Adjusted R^2	0.024	0.023	0.029	0.017	0.012	0.042

The table shows the OLS regressions for which the dependent variable is the three-day cumulative abnormal return estimated from market model residuals. *p*-values are reported below the coefficients. The sample of successful and unsuccessful acquisitions by publicly listed US acquirers is from the SDC Merger and Acquisition Database for the period 1980–2002. It includes all deals involving U.S. private targets using 100% equity payment or public targets with 100% cash or 100% equity payment. The standard deviation (std) of the long-term earnings growth forecasts (LTG) requires three or more analysts and is reported in percent and is from I/B/E/S in the month prior to the deal. Top decile LTG (std) is equal to one if the acquirer's std of its long-term earnings growth forecasts is in the top decile of stds among all acquirers with the same number of analysts. Breadth of ownership is defined for the acquirer as the fraction of mutual funds who own the stock in the quarter prior to the acquisition. Mutual fund holdings are calculated as the aggregate mutual fund holdings divided by the total shares outstanding on CRSP in the quarter prior to the acquisition. The relative transaction value (TVMVE) is the total value of consideration paid by the acquirer, excluding fees and expenses, as reported by SDC divided by the market value of equity. Except for model (1) of Panel A, all regression models include, but do not report the control variables from Moeller, Schlingemann and Stulz (2004). In addition to these control variables, we use a dummy variable equal to one if the closest reported Governance Index from Gompers, Ishii, and Metrick (2003) for the acquirer is above the sample median and a dummy variable equal to 1 if the deal is announced during the calendar years 1998–2000. The superscripts [a], [b], and [c] denote statistical significance of the coefficients at the 1, 5, and 10% levels, based on heteroscedasticity-adjusted standard errors. In Panel B, in models (1)–(3) and (4)–(6), * denotes a significant difference, at the 10% level or better, of the coefficient relative to the same coefficient in Panel A for models (2)–(4).

the findings of Masulis, Wang, and Xie (2007) that acquirer returns are higher for firms with better governance, we use a dummy variable equal to one if the Gompers, Ishii, and Metrick (2003) governance index of the acquirer is above the median. Further, to make sure that our results are not due to the acquisition wave of the late 1990s, we introduce a dummy variable that equals one for offers during the period 1998–2000. The variables used in the literature to explain abnormal returns capture a broad range of determinants of these returns, including variables that could proxy for diversity of opinion. Using these variables, therefore, poses a stringent test for the ability of proxies for diversity of opinion to explain abnormal returns. Note first that the measure of dispersion of analyst forecasts is still significant. The coefficient is now slightly smaller in absolute value, so that the impact of a two-standard deviation change in diversity of opinion on abnormal returns is 2.6%. It follows that the significance of the coefficient on dispersion of analyst forecasts is not due to this variable serving as a proxy for other variables used to explain acquirer abnormal returns. Regression (3) in Panel A shows that our alternate measure of diversity of opinion constructed from analyst forecasts, which is a dummy variable for firms in the top decile of analyst forecasts dispersion given their number of analyst forecasts, has a negative significant coefficient for acquisitions of public firms paid for with equity.

Regression (4) in Panel A uses breadth of ownership, defined as the fraction of mutual funds that own a stock. We use this variable as an alternative measure of the slope of the demand curve for shares. In the theoretical model of Chen, Hong, and Stein (2002), breadth of ownership is negatively correlated with diversity of opinion. In their model, there is less breadth of ownership when there are more pessimistic investors who would like to sell short but cannot. Because mutual fund holdings become more important during the sample period, and since breadth of ownership is correlated with mutual fund holdings, we follow Chen, Hong, and Stein (2002) and control for aggregate mutual fund holdings. The breadth-of-ownership variable is available for a much larger number of acquisitions than our proxies for diversity of opinion derived from analyst forecasts. Though we report the regression for the larger sample, the results are similar for the smaller sample of acquisitions for which our proxies derived from analyst forecasts are available. We find that breadth of ownership is positive and significant as predicted.

We would expect that acquisitions that are larger relative to the equity capitalization of the acquirer have lower abnormal returns when they increase the float. We find that the coefficient on the relative size variable (i.e., the value of the consideration divided by the equity capitalization of the bidder) is significantly negative as expected in regressions (2)–(4). Furthermore, the economic significance of the coefficient is substantial. The average abnormal return for equity offers for public firms in our sample with analyst forecasts is −2.8%. An increase of one standard deviation in the relative size of the acquisition decreases abnormal returns by 1%, so that the abnormal return becomes −3.8%.

For regression (5), we add to regression (2) an interaction of our diversity-of-opinion proxy with the relative size of the offer. We expect this variable to have a negative significant coefficient, yet it does not. However, when we add the interaction variable, neither relative size nor our diversity-of-opinion proxy is significant, as is shown in regression (5). If we omit these two variables, the interaction variable has a significant negative coefficient (not shown), suggesting that multicollinearity could explain the lack of success of the interaction in regression (5). Regression (6) uses breadth of ownership instead of the standard deviation of analyst forecasts. The interaction of relative size with breadth of ownership is not significant either.

3.2 Testing the diversity-of-opinion models using target data

Baker, Coval, and Stein (2006) predict that acquirer abnormal returns in a stock acquisition fall as the proportion of sleepy investors in the target falls. They assume that individual investors are more likely to be sleepy investors than institutional investors, so the individuals hold on to the shares of the acquirer they receive without much thought. Investors who are not sleepy are assumed to have a low opinion of the acquirer; so they want to sell the shares received from the acquirer. Therefore, shares received by target institutional investors are more likely to contribute to the float and hence, with a downward-sloping demand for shares, to lead to a drop in the acquirer's share price. They relate acquirer abnormal returns to the proportion of target shares held by institutional investors and find, as expected, that acquirer abnormal returns fall as this proportion increases. This proportion is calculated using the shares held by institutional investors during the quarter prior to the takeover announcement divided by the total number of shares at the end of the same quarter, both collected from the Thomson Financial CDA/Spectrum database. In a regression we do not report, we find that the proportion of institutional investors in the target stock has a negative significant coefficient and the dispersion of analyst forecasts measure remains significant. Regression (7) of Panel A of Table 3 adds the fraction of target shares held by institutions as an additional interacting variable to the interaction of diversity of opinion and the relative size of the offer. The triple interaction of the diversity-of-opinion proxy, the relative size of the offer, and ownership by institutions of target shares has a positive insignificant coefficient when we use the standard deviation of analyst forecasts as shown in regression (7) and the top decile dummy variable proxy (not shown). In these regressions, the diversity-of-opinion proxy has a negative significant coefficient, but the interaction does not. The last regression, model (8), uses a triple interaction with breadth of ownership, the relative size of the offer, and the ownership of the target shares by institutions. In that regression, the triple interaction is significant but breadth of ownership is not.

We estimate regressions using other target characteristics. In doing so, we face the problem that requiring analyst data for the target as well as the bidder shrinks our dataset to less than one-third its size in Table 3. In contrast, using breadth of ownership has little impact on our sample size. We therefore reestimate regression (4) of Table 3 of Panel A by adding breadth of ownership of the target and the target run-up. We include the target run-up to capture a possible capital gains lock-in effect. We would expect target shareholders to be less likely to sell the shares received from the acquirer if their tax basis is lower, so the float increase would be smaller. As pointed out by Baker, Coval, and Stein (2006), capital gains may also increase the premium paid by the acquirer if the target shareholders expect to sell their shares because they do not want to hold the shares of the acquirer. If this latter effect dominates the former effect, capital gains could worsen the abnormal return of the acquirer. With the capital lock-in effect we would expect the target run-up to be significantly positive for acquisitions of public firms paid for with equity but not for acquisitions of public firms paid for with cash. Our results (not reported) are robust to the inclusion of the target run-up. The target run-up is significant and positive, which is consistent with a capital gains lock-in effect. Breadth of ownership of the target has a positive insignificant coefficient that is not significantly different from the coefficient on breadth of ownership of the acquirer, which continues to have a positive significant coefficient.

3.3 Acquirer returns and diversity of opinion for private firm acquisitions and acquisitions of public firms paid for with cash

Panel B of Table 3 estimates regressions (2) through (4) of Panel A for acquisitions of private firms paid for with equity and for acquisitions of public firms paid for with cash. For the diversity-of-opinion proxies, an asterisk indicates that the coefficient on the proxy is significantly different from the coefficient on the same proxy in the regressions of Panel A at the 10% level or better.[10] All the coefficients on proxies for diversity of opinion (except for one) are significantly different from their values in Panel A and, as expected, the proxies for diversity of opinion do not have significant coefficients. For acquisitions of public firms paid for with cash, the proxies constructed from analyst forecasts do not have significant coefficients. However, breadth of ownership has a significant coefficient that has the opposite sign from its sign in the regression for equity offers for public

[10] To evaluate the significance of the difference, we estimate a pooled regression in which we allow the intercept and slopes of all variables except the industry dummy to depend on the type of transaction.

firms. This result is inconsistent with the predictions of the diversity-of-opinion models. We also estimate regressions with target characteristics (not reported). Strikingly, when we consider acquisitions of public firms paid for with cash, the coefficient on target run-up is very similar, though insignificant, to the coefficient on target run-up for acquisitions paid for with equity even though the target run-up is not expected to affect the tendering decision in cash offers.[11] This result is therefore not supportive of the diversity-of-opinion models.

With the diversity-of-opinion models, we would not expect the relative size variable to have a negative significant coefficient for acquisitions of private firms and for acquisitions of public firms paid for with cash. In Panel B of Table 3, the coefficient on relative size is positive but never significant.

The results presented thus far show that diversity-of-opinion proxies are related to acquirer returns in a way consistent with our assumptions about the changes in float. We assume that cash offers do not increase the float, and equity offers for public firms increase the float substantially more than equity offers for private firms. To check whether these assumptions are reasonable, we investigate the changes in trading volume around acquisitions since we cannot measure the float directly. If public firm acquisitions paid for with equity increase the float more than private firm acquisitions paid for with equity, we expect a greater increase in shares traded after the acquisition for acquisitions of public firms than for acquisitions of private firms paid for with equity.

To investigate this, we collect data on the acquirer's stock trading volume from CRSP for two windows of 50 days. The first window ends 25 days before the announcement and the second window starts 25 days after completion. The median percentage change in volume for public acquisitions paid for with equity is 58.33%; in contrast, it is 29.36% for acquisitions of private firms paid for with equity and 13.07% for acquisitions of public firms paid for with cash. The median changes are significantly different from zero and from each other. With our assumption that the float increases only with acquisitions of public firms paid for with equity, the size of an acquisition of a private firm paid for with equity or of a public firm paid for with cash should have a weaker relation with the volume percentage change. To test this, we regress the volume percentage change on a constant, a dummy for private equity acquisitions (PrivEq), a dummy for public firm acquisitions paid for with cash (PubCash), a dummy for private firm acquisitions paid for with cash (PrivCash), the ratio of deal value to bidder market capitalization (TVMVE), interactions of the

[11] A number of papers predict that the supply of shares in response to a tender offer for cash is upward sloping because the realization of capital gains decreases the gain from tendering shares. Anderson and Dyl (2004) review this literature and provide evidence of this effect for self-tender offers.

dummy variables with the ratio of deal value bidder market capitalization, and year dummies. The estimated regression is:

$$\begin{aligned}\text{Volume percentage change} = {} & 0.6813 + 0.0291 \times \text{PrivEq} \\ & - 0.2696 \times \text{PubCash} + 0.9134 \times \text{TVMVE} \\ & - 0.8133 \times [\text{PrivEq} \times \text{TVMVE}] \\ & - 0.9201 \times [\text{PubCash} \times \text{TVMVE}] \qquad (2)\end{aligned}$$

All coefficients are significant at the 1% level except for the coefficient on PrivEq. The regression is estimated on 1248 observations and has an adjusted R^2 of 0.070. The impact of TVMVE on the volume percentage change (i.e., 0.9134–0.8133) is not significantly different from zero for acquisitions of private firms paid for with equity or acquisitions of public firms paid for with cash. It is therefore reasonable to believe that acquisitions of private firms paid for with equity have a negligible effect on the bidder's float, and so our result that there is no relation between abnormal returns for such acquisitions and dispersion of analyst forecasts is consistent with the diversity-of-opinion models.

In summary, we find results that are generally consistent with diversity-of-opinion models: (i) acquirer announcement returns for equity offers for public firms decrease as diversity of opinion increases but announcement returns for other offer types do not, and (ii) larger equity offers for public firms have worse acquirer abnormal returns, while the size of the offer does not matter for other offers. Consistent with the diversity-of-opinion models, we also find that interacting breadth of ownership with the relative size of the offer and target institutional ownership is helpful in explaining acquirer returns.

4. Information Asymmetry

Our proxies for diversity of opinion help explain the acquirer abnormal returns for acquisitions of public firms paid for with equity as predicted by the diversity-of-opinion models, but we also find results inconsistent with these models. In particular, we find that an increase in breadth of ownership leads to significantly higher returns for cash offers for public companies. Such a result could be understood if breadth-of-ownership proxies for information asymmetry since, in the presence of information asymmetry, a cash offer is good news about the value of the bidder's common stock. To understand better the extent to which our results can be explained by asymmetric information models, we investigate the relation between abnormal returns and the proxies for information asymmetry used in Dierkens (1991) discussed earlier.

Table 4 presents the regression results explaining acquirer returns with several proxies for asymmetric information added to the dispersion of long-term growth forecasts and control variables (not reported) from Table 3. Regression (1) of Table 4 adds idiosyncratic volatility to regression (2) of Table 3. We find that idiosyncratic volatility is highly significant with a negative coefficient. The economic significance of idiosyncratic volatility is substantial. The standard deviation of idiosyncratic volatility is 1.56%. Since the regression coefficient on idiosyncratic volatility in regression (1) is −0.9560, a one-standard deviation increase in idiosyncratic volatility corresponds to a decrease in the abnormal return of 1.49%. Since the average abnormal return for equity acquisitions of public firms is −2.834%, an acquisition by an acquirer with an idiosyncratic volatility of one standard deviation above the mean has an abnormal return of −4.42%. When idiosyncratic volatility is added to the regression, dispersion of analyst forecasts is no longer significant. We then use our alternative dummy variable measure of high dispersion of analyst forecasts in regression (2). That measure has an insignificant coefficient also when idiosyncratic volatility is added. In regression (3), we add the standard deviation of earnings announcements to regression (1). We still find that idiosyncratic volatility has a negative significant coefficient. The standard deviation of earnings announcement returns is not significant. Finally, in regression (4) we add breadth of ownership and the mutual fund holding variable to regression (1). We find that breadth of ownership is not significant. It follows that idiosyncratic volatility dominates our proxies for diversity of opinion. However, several recent papers use idiosyncratic volatility as a proxy for diversity of opinion (see Gebhardt, Lee, and Swaminathan (2001); Danielsen and Sorescu, (2001) Diether, Malloy, and Scherbina (2002); Boehme, Danielsen, and Sorescu (2006)). Since it is possible that idiosyncratic volatility is a better measure of diversity of opinion than dispersion of analyst forecasts, these results might be consistent with the diversity-of-opinion models and show that the diversity of opinion proxies used in Table 3 are noisy.

Regressions (5)–(8) reestimate models (1)–(4) on the sample of acquisitions of private firms paid for with equity. The coefficient on idiosyncratic volatility is positive but insignificant in these regressions. In Table 2 we show that the average and median abnormal returns for the announcement of acquisitions of private firms paid for with equity are significantly positive. Consequently, the lack of significance on the coefficient of idiosyncratic volatility is inconsistent with the view that for a given distribution of returns, the expected return conditional on good news increases with uncertainty. Except for regressions (6) and (7), the coefficient on idiosyncratic volatility is significantly higher than in the regressions for acquisitions of public firms paid for with equity. Surprisingly, the volatility of the earnings residual is significant and positive in regression (7).

Table 4
Cross-sectional regressions with information asymmetry measures

	Public targets | Equity				Private targets | Equity				Public targets | Cash			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
LTG (std)	−0.0006 0.681		−0.0003 0.807	−0.0006 0.692	−0.0010 0.369		−0.0011 0.359	−0.0010 0.406	0.0021 0.139		0.0018 0.270	0.0031^{b} 0.043
Top decile LTG (std)		−0.0208 0.178				0.0047 0.714				0.0306^{c*} 0.078		
Volatility	-0.9560^{b} 0.011	-0.8775^{b} 0.012	-0.8742^{c} 0.060	-0.8500^{b} 0.030	0.2340^{*} 0.595	0.0645 0.876	−0.1383 0.795	0.3019^{*} 0.523	0.8323^{c*} 0.063	0.8873^{c*} 0.041	1.0986^{b*} 0.042	0.4812^{*} 0.337
Earnings residual (std)			−0.0798 0.528				0.3066^{c*} 0.090				−0.1776 0.267	
Breadth of ownership				0.0478 0.476				−0.1283 0.506				0.0315 0.695
Mutual fund holdings				−0.1261 0.114				−0.0310 0.825				0.1329^{*} 0.161
TVMVE	-0.0272^{a} 0.002	-0.0270^{a} 0.002	-0.0260^{a} 0.004	-0.0267^{a} 0.005	0.0096 0.395	0.0107 0.343	0.0113 0.294	0.0090 0.428	0.0088 0.444	0.0087 0.463	0.0198^{a} 0.006	0.0117 0.297
Intercept	0.0086 0.719	0.0051 0.828	0.0090 0.705	0.0738^{c} 0.078	0.0388^{*} 0.123	0.0202^{*} 0.381	0.0383 0.152	0.0479 0.122	-0.0459^{c} 0.047	−0.0341 0.151	-0.0458^{c} 0.050	-0.0509^{b} 0.030
Observations	682	682	670	665	551	551	524	538	211	211	192	196
Adjusted R^2	0.076	0.080	0.070	0.073	0.026	0.024	0.030	0.024	0.059	0.065	0.080	0.059

The table gives the OLS regressions for which the dependent variable is the three-day cumulative abnormal return estimated from market model residuals. *p*-values are reported below the coefficients. The sample of successful and unsuccessful acquisitions by publicly listed US acquirers is from the SDC Merger and Acquisition Database for the period 1980–2002. It includes all deals involving US private targets using 100% equity payment or public targets with 100% cash or 100% equity payment. The standard deviation (std) of the long-term earnings growth forecasts (LTG) uses I/B/E/S data from the month prior to the announcement and is reported in percent. Top decile LTG (std) is equal to one if the acquirer's standard deviation of its long-term earnings growth forecasts is in the top decile of standard deviations among all acquirers with the same number of analysts. Volatility is the standard deviation of the market-adjusted residuals of the daily stock returns measured during the period from 205 to six days prior to the acquisition announcement. Earnings Residual (std) is the standard deviation of all three-day cumulative abnormal returns around earnings announcements from I/B/E/S using the market model over the five-year period preceding the acquisition announcement. Breadth of ownership is the fraction of mutual funds that own the acquirer's stock in the quarter prior to the acquisition. Mutual fund holdings are calculated as the aggregate mutual fund holdings divided by the total shares outstanding on CRSP in the quarter prior to the acquisition. The relative transaction value (TVMVE) represents the total value of consideration paid by the acquirer, excluding fees and expenses as reported by SDC divided by the market value of equity. Except for model (1) all regression models include, but do not report the control variables as defined in Table 3. Superscripts ^{a}, ^{b}, and ^{c} denote statistical significance of the coefficients at the 1, 5, and 10%, respectively, based on heteroscedasticity-adjusted standard errors. In models (5)–(8) and (9)–(12), * denotes a significant difference at the 10% level or better of the coefficient relative to the same coefficient in models (1)–(4).

The models that focus on diversity of opinion predict that the announcement return of an acquisition paid for with cash should not be related to a proxy for diversity of opinion since no new equity is issued. In contrast, information asymmetry models imply that the announcement return on acquisitions paid for with cash should increase with variables that measure information asymmetry. Not issuing equity when financing an acquisition must be interpreted as positive news about the value of equity and it must be more positive news when there is greater uncertainty about the true value of the firm. Regressions (9)–(12) use the sample of acquisitions of public firms paid for with cash. For these regressions, we find that the coefficient on idiosyncratic volatility is positive and significant. The absolute value of the coefficient is similar to the absolute value of the coefficient for the acquisitions of public firms paid for with equity. This result is inconsistent with the predictions of diversity-of-opinion models. It is, however, fully consistent with the information asymmetry models. Note that this result cannot be explained by the prediction that the expected return conditional on good news increases with volatility since the average and median abnormal returns for the cash offers in our sample are negative.

The Myers and Majluf (1984) model implies that an equity issue conveys information about the value of assets in place. Jain (1992) tests that implication of the model by regressing changes in analyst forecasts around the time of an equity issuance on the issue's abnormal return. Though we do not report the results in a table, we compare changes in forecasts for the three types of acquisitions we focus on. Using median changes, we find acquisitions of private firms paid for with equity have the highest change, public acquisitions paid for with cash have the lowest change, and acquisitions of public firms paid for with equity are the middle. The only significant difference using the long-term earnings growth forecasts is between acquisitions of private firms paid for with equity and acquisitions of public firms paid for with cash. The change in yearly earnings forecasts normalized by the stock price is significantly higher for private firm acquisitions paid for with equity than for public firm acquisitions paid for with cash. However, the difference between public firm acquisitions paid for with equity and public firm acquisitions paid for with cash is not significant. In sum, changes in analyst forecasts are not supportive of the Myers and Majluf (1984) prediction, since with that prediction the change should be lowest for acquisitions of public firms paid for with equity. Of course, this conclusion is based on the assumption that there are no systematic differences in synergies across acquisition types. For the evidence to be reconciled with Myers and Majluf, we would need to believe that somehow equity offers have greater synergy than cash offers. There is no basis in the literature to conclude this.

In summary, we find that proxies for information asymmetry help explain acquirer abnormal returns as predicted by the information asymmetry models. Perhaps most importantly, we find that the announcement return for cash offers for public firms increases with acquirer stock return volatility, a result that is inconsistent with diversity-of-opinion models if stock return volatility also proxies for diversity of opinion. However, subsequent changes in analyst forecasts do not seem supportive of the information asymmetry models.

5. Changes in Forecast Dispersion and Bidder Abnormal Returns

The information asymmetry models imply that investors learn from corporate actions, so that information asymmetry should fall following a merger announcement. Table 2 shows that the mean, but not the median, change in dispersion of analyst forecasts from before the announcement of an acquisition to after the announcement is significantly negative. This change in the dispersion of analyst forecasts makes it possible for the dispersion of analyst forecasts to proxy for resolution of uncertainty. The models of Pástor and Veronesi (2006) and Johnson (2004) predict a positive relation between a firm's value and uncertainty about its long-run expected growth. We proxy for uncertainty regarding a firm's long-run expected growth with the dispersion of long-term growth analyst forecasts. We first compare the abnormal returns for bidders, where the change in dispersion is negative with the abnormal return for bidders where the change is positive, ignoring the firms where the change is zero. The bidders with a decrease in dispersion of analyst forecasts have lower abnormal returns by 1.26% on average for acquisitions of public firms paid for with equity. After the beginning of 1998 the difference is a relatively large 3%. In contrast, for private firms paid for with equity, bidders experience abnormal returns that are 1.22% higher when the dispersion of analyst forecasts decreases as opposed to increases during the full sample period. While the former result is consistent with resolution-of-uncertainty models, the latter is not.

In Panel A of Table 5, we add to regression (2) of Panel A of Table 3 the dispersion of analyst forecasts in month $t + 1$. The change in dispersion has a correlation of -0.2 with the level of the dispersion of analyst forecasts. Consequently, firms with high dispersion of analyst forecasts are expected to see a drop in their dispersion of analyst forecasts. To account for this, we also add to the regression the median analyst forecast at $t - 1$ and the median analyst forecast at $t + 1$. We see that all four variables are significant. A decrease in dispersion of analyst forecasts is associated with a worse abnormal return as predicted. Further, an increase in analyst forecasts is associated with a better abnormal return. This result is supportive of Myers and Majluf (1984), but it could also mean that analysts become more pessimistic when an offer is poorly received by the market. In regression (2), we add idiosyncratic volatility to model (1). Its coefficient is significantly negative. Regression (2) demonstrates that

Table 5
Cross-sectional regressions of abnormal returns on pre- and post-forecasts and control variables

	Public equity (1)	Public equity (2)	Private equity (3)	Private equity (4)	Public cash (5)	Public cash (6)
LTG (std)$_{month-1}$	-0.0065^{b}	-0.0036^{b}	-0.0004^{*}	-0.0008	0.0048^{*}	0.0029^{*}
	0.033	0.046	0.819	0.710	0.073	0.266
LTG (std)$_{month+1}$	0.0061^{b}	0.0040^{b}	-0.0010^{*}	-0.0011^{*}	-0.0035^{*}	-0.0014^{*}
	0.037	0.017	0.604	0.623	0.226	0.596
LTG (median)$_{month-1}$	-0.0055^{a}	-0.0047^{a}	-0.0006^{*}	-0.0010^{*}	-0.0034	0.0006
	0.004	0.009	0.669	0.484	0.408	0.891
LTG (median)$_{month+1}$	0.0043^{b}	0.0040^{b}	0.0012	0.0014	0.0045	-0.0004
	0.024	0.029	0.414	0.357	0.284	0.912
Volatility		-0.8469^{b}		0.4442^{*}		0.8652^{c*}
		0.055		0.444		0.054
TVMVE	-0.0253^{a}	-0.0275^{a}	0.0489^{a*}	0.0435^{b*}	0.0079^{*}	0.0102^{*}
	0.007	0.003	0.009	0.024	0.503	0.409
Intercept	0.0658^{c}	0.0657	0.0176	0.0186	-0.0422^{c}	-0.0428^{c}
	0.069	0.115	0.471	0.488	0.097	0.085
Observations	707	673	574	531	211	204
Adjusted R^2	0.101	0.091	0.020	0.022	0.034	0.061

The table shows the OLS regressions where the dependent variable is the three-day cumulative abnormal return measured using market model residuals. *p*-values are reported below the coefficients. The sample of successful and unsuccessful acquisitions by publicly listed US acquirers is from the SDC Merger and Acquisition Database for the period 1980–2002. It includes all deals involving US private targets using 100% equity payment or public targets with 100% cash or 100% equity payment. The standard deviation (std) of the long-term earnings growth forecasts requires three or more analysts and is reported in percent. The std and median of the long-term growth forecasts are measured in the month before the announcement and in the month after the announcement. Volatility is the std of the market-adjusted residuals of the daily stock returns measured during the period from 205 days to six days prior to the acquisition announcement. The relative transaction value (TVMVE) represents the total value of consideration paid by the acquirer, excluding fees and expenses, as reported by SDC divided by the market value of equity. All regression models include, but do not report, the control variables as defined in Table 3. Superscripts ^{a}, ^{b}, and ^{c} denote statistical significance of the coefficients at the 1, 5, and 10%, based on heteroscedasticity-adjusted standard errors. In models (3)–(4) and (5)–(6), * denotes a significant difference, at the 10% level or better, of the coefficient relative to the same coefficient in models (1) and (2).

models based on resolution of uncertainty are helpful in understanding bidder abnormal returns for acquisitions of public firms paid for with equity, but when resolution of uncertainty is taken into account, the coefficient on idiosyncratic volatility is still significant.

Regressions (3) and (4) show regression estimates for private firm acquisitions paid for with equity. The coefficients on the resolution-of-uncertainty variables are not significant, but the coefficient on idiosyncratic volatility is positive and significant. Finally, regressions (5) and (6) show estimates for acquisitions of public firms paid for with cash. The coefficients on the resolution-of-uncertainty variables have the wrong sign, are insignificant, and are significantly different from regressions (1) and (2). The resolution-of-uncertainty models predict that resolution of uncertainty is associated with lower abnormal returns for all types of offers. Consequently, the evidence in regressions (3)–(6) is not supportive of these models.[12]

6. Understanding Differences in Acquirer Abnormal Returns Across Offer Types

The existence of large differences in average abnormal returns among the three types of acquisitions we focus on (shown in Table 2 for our sample) are well known (see, for instance, the references in Moeller, Schlingemann, and Stulz (2004)). In this section, we investigate whether these differences in abnormal returns across offer types can be understood using our proxies for diversity of opinion, information asymmetry, and resolution of uncertainty.

To examine whether models that focus on diversity of opinion can explain these differences, we proceed as follows. We first estimate a pooled regression with all acquisitions for which we have proxies for diversity of opinion derived from analyst forecasts. We regress the abnormal returns on an intercept, a dummy variable for cash offers of public targets, and a dummy variable for equity offers of private targets. Regression (1) in Table 6 shows that the intercept is significantly negative and each of the dummy variables is significantly positive, so that the abnormal returns for different offer types are significantly different from each other. We then expand the regression by adding the dispersion of analyst forecasts and interactions of the dispersion of analyst forecasts with the two dummy variables. Regression (2) shows that private firm acquisitions still have a significantly unconditional higher abnormal return than public firm acquisitions paid for with equity (difference of 3.85% with a p-value <0.001). However, the cash acquisitions of public firms do not have an unconditionally higher abnormal return than the equity acquisitions of public firms. While cash offers have a higher abnormal return of 2.52% with a p-value smaller than 0.001 when not controlling for the dispersion of analyst forecasts, the difference falls to 0.69% with a p-value of 0.357 when doing so. An even more dramatic change occurs when controlling for breadth of ownership since the difference falls from 2.64 to -1.06% (not reported). Finally, regression (3) adds the control variables. Again, the private firm acquisition dummy variable is significantly positive, but the dummy for public firm acquisitions paid for with cash is not. In summary,

[12] A concern with our results is that the post-announcement forecasts might be difficult to interpret because it is not completely clear how analysts take into account the impact of the acquisition on long-term growth if the merger has not been completed when these forecasts are made. From conversations with personnel at I/B/E/S and sell-side analysts, we understand that forecasts do not incorporate the specific acquisition if the offer is not completed but may take into account the general implications of the offer for the firm as a stand-alone entity. To ensure this issue is not driving our results, we rerun the test with a sample of offers where the merger completion predates the publication of the post-announcement forecasts. We find similar, but stronger, results.

Table 6
Cross-sectional regressions of abnormal returns using payment and target type indicators

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
(Private | Equity) indicator	0.0475[a]	0.0385[a]	0.0248[a]	0.0534[a]	−0.0059	−0.0117	0.0469[a]	0.0417[a]	0.0239[a]
	0.000	0.000	0.000	0.000	0.585	0.288	0.000	0.000	0.008
(Public | Cash) indicator	0.0252[a]	0.0069	−0.0044	0.0274[a]	−0.001	−0.0101	−0.0247[a]	0.0095	−0.0028
	0.000	0.357	0.622	0.000	0.907	0.249	0.000	0.196	0.759
LTG $(std)_{month-1}$		−0.0029[b]	−0.0030[b]					−0.0063[b]	−0.0071[a]
		0.012	0.012					0.018	0.005
LTG $(std)_{month-1}$ × (Private | Equity)		0.0023	0.0028[c]					0.0060[c]	0.0068[b]
		0.100	0.055					0.060	0.029
LTG $(std)_{month-1}$ × (Public | Cash)		0.0047[b]	0.0050[b]					0.0103[b]	0.0105[a]
		0.020	0.011					0.011	0.006
Volatility					−0.2172	−0.0289			
					0.376	0.916			
Volatility × (Private | Equity)					1.4590[a]	1.3128[a]			
					0.000	0.000			
Volatility × (Public | Cash)					1.0787[a]	0.7522[b]			
					0.002	0.030			
LTG $(std)_{month+1}$								0.0042	0.0047[c]
								0.102	0.066
LTG $(std)_{month+1}$ × (Private | Equity)								−0.0039	−0.0044
								0.209	0.157
LTG $(std)_{month+1}$ × (Public | Cash)								−0.0067	−0.0066[c]
								0.101	0.093

TVMVE			−0.0202[b]			−0.0216[a]			−0.0209[b]
			0.027			0.000			0.032
TVMVE × (Private | Equity)			0.0339[b]			0.0223[a]			0.0663[a]
			0.033			0.000			0.001
TVMVE × (Public | Cash)			0.0206			0.0210[a]			0.0203
			0.115			0.003			0.144
Intercept	−0.0286[a]	−0.0164[a]	−0.0385[a]	−0.0229[a]	−0.0165[a]	−0.0187	−0.0289[a]	−0.0137[b]	−0.0448[a]
	0.000	0.001	0.002	0.000	0.007	0.215	0.000	0.014	0.001
Observations	1545	1545	1538	3880	3880	3844	1499	1499	1492
Adjusted R^2	0.052	0.059	0.078	0.043	0.073	0.098	0.053	0.067	0.081

The table gives the OLS regressions where the dependent variable is the three-day cumulative abnormal return measured using market model residuals. *p*-values are reported below the coefficients. The sample of successful and unsuccessful acquisitions by publicly listed US acquirers is from the SDC Merger and Acquisition Database for the period 1980–2002. it includes all deals involving US private targets using 100% equity payment (Private | Equity) or public targets with 100% cash (Public | Cash) or 100% equity payment (Public | Equity). The standard deviation (std) of the long-term earnings growth forecasts requires three or more analysts and is reported in percent. LTG (std) $_{\text{month+1}}$ is measured after the announcement and LTG (std) $_{\text{month−1}}$ is measured before the announcement. Volatility is the std of the market-adjusted residuals of the daily stock returns measured during the period starting from 205 days to six days prior to the acquisition announcement. The relative transaction value (TVMVE) represents the total value of consideration paid by the acquirer, excluding fees and expenses as reported by SDC divided by the market value of equity. Regression models (3), (6), and (9) include, but do not report the control variables as defined in Table 3. Model (7) restricts the sample of model (1) to further include data on LTG for the month after the announcement. Models (8) and (9) include, but do not report, the median of the long-term earnings growth forecasts measured in the month before and after the announcement. Respectively, Superscripts a, b, and c denote statistical significance of the coefficients at the 1, 5, and 10%, based on heteroscedasticity-adjusted standard errors.

diversity-of-opinion variables can explain the difference between cash and equity offers for public firms, but not between equity offers for public firms and equity offers for private firms.

We turn next to the ability of information asymmetry to explain differences in abnormal returns across offer types. We are able to use a larger sample when we do not use proxies for diversity of opinion derived from analyst forecasts. In models (4) and (5), we proceed as we did with the dispersion of analyst forecasts and the breadth of ownership and use the largest possible sample of acquisitions. We find that when excluding our usual control variables and without controlling for idiosyncratic volatility, the average abnormal return of cash offers for public firms exceeds the average abnormal return of equity offers for public firms by 2.74% with a p-value of less than 0.001. Taking into account idiosyncratic volatility, this difference falls to −0.10% with a p-value of 0.907. While the dispersion of analyst forecasts and breadth of ownership fail to explain the higher abnormal return of private firm acquisitions paid for with equity, idiosyncratic volatility seems successful in doing so. Without taking into account idiosyncratic volatility, private firm acquisitions have a higher abnormal return than public firm acquisitions paid for with equity of 5.34% with a p-value of less than 0.001. This difference falls to −0.59% with a p-value of 0.585 when we take into account idiosyncratic volatility. In other words, if idiosyncratic volatility is set equal to zero, there are no significant differences between the abnormal returns of the three types of acquisitions. The last regression uses the interacted control variables and shows that our result continues to hold.

Finally, we consider the resolution-of-uncertainty models for completeness. We already found that we can reject these models. Here, we find that they cannot explain the differences in abnormal returns across offer types. As can be seen in Table 6, in regressions (7)–(9) the unconditional abnormal return for announcements of acquisitions of private firms paid for with equity is positive and significant when we take into account resolution of uncertainty and our control variables.

7. Additional Robustness Checks of the Results and Their Interpretation

7.1 Do diversity of opinion and information asymmetry have a permanent impact?

An additional prediction coming out of the models in this article is that the effects documented should be permanent. If the effects we document are temporary, acquisitions of public firms by acquirers with high valuations and high dispersion of analyst forecasts should have a higher abnormal return after the acquisition announcement of similar acquisitions by acquirers with high valuations and low dispersion of analyst forecasts. We find no evidence consistent with this prediction when we use abnormal returns for the 20 days following the offer announcement.

Another approach to evaluate whether the effect we document is transitory is to estimate a regression of postannouncement returns on announcement returns, dispersion of analyst forecasts, and dispersion of analyst forecasts interacted with announcement returns. Reversal of the announcement return would imply a negative coefficient on the announcement return. Reversal of the effect we document requires a negative coefficient on the interaction of dispersion of analyst forecasts and the announcement return. When we estimate this regression (not reported) for offers to acquire public firms financed with equity, the coefficient on the announcement return is 0.0220 with a p-value of 0.064, the coefficient on the interaction is -0.0164 with a p-value of 0.399, and the coefficient on dispersion of analyst forecasts is -0.0040 with a p-value of 0.168. Since the coefficient on the interaction is not significant, we find no evidence of reversal for equity offers for public firms. Surprisingly, though, when we estimate a similar regression for equity offers for private firms, the coefficient on the interaction is negative and significant.

7.2 Do the results hold for alternate definitions of returns?

The abnormal returns used so far are estimated over a short window around the announcement date. Offers often get revised, so that information about the offer that we use in our regressions may become available later.[13] Authors have also argued that the information revealed by the acquisition announcement does not necessarily get impounded in the stock price all at once. We therefore estimate our regressions using a different definition of abnormal returns, the buy-and-hold returns from the day before the announcement to the effective date. Our results hold for this alternate definition of returns.

7.3 Does diversity of analyst forecasts proxy for changes in analyst forecasts?

Scherbina (2003) provides evidence that analyst forecasts dispersion is related to the prior performance of the firm. This raises the concern that our measure of forecast dispersion could proxy for changes in the level of forecasts or for the level of forecasts. To examine this possibility, we estimate, but do not report, regressions where we add changes in the long-term forecast as well as the level of the forecast. We first add the past-year change in the long-term growth forecast to regression (2) from Table 3, which includes the forecast dispersion and prior control variables. The forecast dispersion remains significant with no change in the coefficient, while the past-year change is not significant. Finally, we add to that regression the quarterly earnings surprise because Scherbina concludes that analysts are too optimistic when dispersion is high. The quarterly

[13] This is only a problem for the multiple regressions in which we control for deal characteristics. None of our key conclusions are sensitive to the use of these controls.

earnings surprise is calculated as the actual quarterly earnings minus the most recent median (mean) analyst forecast if the announcement occurred during the period starting one month before the announcement and ending one month after the announcement. If no announcement occurred in that period, the earnings surprise is zero. Adding the quarterly earnings surprise does not change our results. We repeat these regressions for private firm acquisitions paid for with equity. The coefficient on the dispersion of long-term forecasts is insignificant in each model for private firm acquisitions.

8. Conclusion

Theoretical papers predict that acquirer abnormal returns should be related to proxies for diversity of opinion among investors and information asymmetry. We find that abnormal returns for equity offers for public firms decrease as these proxies increase, whereas abnormal returns for other offer types are either unaffected or increase.

Though some of our results can be explained by both diversity-of-opinion models and information asymmetry models, and the evidence is not completely consistent with any of the models, we uncover some results that can be explained only by models that emphasize information asymmetry. This suggests that these models may offer a more complete account of the overall evidence than diversity-of-opinion models. This is specifically the case for our finding that the abnormal return associated with acquisitions of public firms paid for with cash significantly increases with some measures of information asymmetry or diversity of opinion.

It is well known that the three types of acquisitions we focus on in this article have significantly different average abnormal returns. The existing literature has not succeeded in explaining these differences. We find that the dispersion of analyst forecasts and the breadth of ownership are successful in explaining the difference in abnormal returns between cash and equity acquisitions of public firms but not between equity acquisitions of public firms and of private firms. Strikingly, however, differences in abnormal returns between acquisition types can be explained by differences in the relation between abnormal returns and idiosyncratic volatility across offer types.

References

Anderson, A. M., and E. Dyl. 2004. Determinants of Premiums on Self-Tender Offers. *Financial Management* 33:25–45.

Baker, M., J. D. Coval, and J. C. Stein. 2006. Corporate Financing Decisions When Investors Take the Path of Least Resistance. *Journal of Financial Economics* 84:266–98.

Boehme, R.D., B. R. Danielsen, and S. M. Sorescu. 2006. Short Sale Constraints, Differences of Opinion, and Overvaluation. *Journal of Financial and Quantitative Analysis* 41:455–87.

Brown, S. J., and J. B. Warner. 1985. Using Daily Stock Returns, the Case of Event Studies. *Journal of Financial Economics* 14:3–31.

Chen, J., H. Hong, and J. C. Stein. 2002. Breadth of Ownership and Stock Returns. *Journal of Financial Economics* 66:171–205.

Danielsen, B. R., and S. M. Sorescu. 2001. Why Do Option Introductions Depress Stock Prices? A Study of Diminishing Short Sale Constraints. *Journal of Financial and Quantitative Analysis* 36:451–84.

Diamond, D. D., and R. E. Verrecchia. 1987. Constraints on Short-Selling and Asset Price Adjustment to Private Information. *Journal of Financial Economics* 18:277–311.

Dierkens, N. 1991. Information Asymmetry and Equity Issues. *Journal of Financial and Quantitative Analysis* 26:181–200.

Diether, K. B. 2004. Long-Run Event Performance and Opinion Divergence. Working Paper, The Ohio State University.

Diether, K. B., C. J. Malloy, and A. Scherbina. 2002. Differences of Opinion and the Cross Section of Stock Returns. *Journal of Finance* 57:2213–41.

Gebhardt, W. R., C. M. Lee, and B. Swaminathan. 2001. Toward an Implied Cost of Capital. *Journal of Accounting Research* 39:135–76.

Gompers, P. A., J. L. Ishii, and A. Metrick. 2003. Corporate Governance and Equity Prices. *Quarterly Journal of Economics* 118:107–55.

Hertzel, M. G., and R. L. Smith. 1993. Market Discounts and Shareholder Gains for Placing Equity Privately. *Journal of Finance* 48:459–85.

Hong, H., T. Lim, and J. C. Stein. 2000. Bad News Travels Slowly: Size, Analyst Coverage, and the Profitability of Momentum Strategies. *Journal of Finance* 55:265–95.

Hong, H., J. Scheinkman, and W. Xiong. 2006. Asset Float and Speculative Bubbles. *Journal of Finance* 61:1073–117.

Jain, P. C. 1992. Equity Issues and Changes in Expectations of Earnings by Financial Analysts. *Review of Financial Studies* 5:669–83.

Johnson, T. C. 2004. Forecast Dispersion and the Cross-Section of Returns. *Journal of Finance* 59:1957–78.

Jovanovic, B., and S. Braguinsky. 2004. Bidder Discounts and Target Premia in Takeovers. *American Economic Review* 94:46–56.

Krasker, W. C. 1986. Stock Price Movements in Response to Stock Issues Under Asymmetric Information. *Journal of Finance* 41:93–106.

La Porta, R. 1996. Expectations and the Cross-Section of Stock Returns. *Journal of Finance* 51:1715–42.

Masulis, R. W., C. Wang, and F. Xie. 2007. Corporate Governance and Acquirer Returns. *Journal of Finance* 62:1851–89.

McCardle, K. F., and S. Viswanathan. 1994. The Direct Entry versus Takeover Decision and Stock Price Performance around Takeovers. *Journal of Business* 67:1–43.

Miller, E. M. 1977. Risk, Uncertainty, and Divergence of Opinion. *Journal of Finance* 32:1151–68.

Moeller, S. B., F. P. Schlingemann, and R. M. Stulz. 2004. Firm Size and the Gains from Acquisitions. *Journal of Financial Economics* 73:201–28.

Myers, S. C., and N. S. Majluf. 1984. Corporate Financing and Investment Decisions When Firms Have Information That Investors Do Not Have. *Journal of Financial Economics* 13:187–221.

Pástor, L., and P. Veronesi. 2006. Was There a Nasdaq Bubble in the Late 1990s? *Journal of Financial Economics* 81:61–100.

Pound, J. 1988. The Information Effects of Takeover Bids and Resistance. *Journal of Financial Economics* 22:207–27.

Scherbina, A. 2003. Analyst Disagreement, Forecast Bias and Stock Returns. Working Paper, Harvard Business School.

Travlos, N. G. 1987. Corporate Takeover Bids, Methods of Payment, and Bidding Firms' Stock Returns. *Journal of Finance* 42:943–63.