Liquidity, Investor-Level Tax Rates, and Expected Rates of Return

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Abstract

Prior research predicts a positive relation between expected rates of return and investor-level tax rates. We provide new theory that predicts that lower liquidity amplifies and higher liquidity attenuates this positive relation. We empirically test our prediction using the cuts to individual investors' maximum statutory tax rates on dividend income and capital gain income enacted by the Jobs and Growth Tax Relief Reconciliation Act of 2003 (JGTRRA03). Although returns are significantly higher in the four years following than in the four years preceding JGTRRA03, consistent with our prediction we find that the increase in returns is significantly smaller among less liquid firms. This result holds for both dividend-paying and non-dividend-paying firms. In addition to offering a new cross-sectional prediction related to tax capitalization, this paper has important implications for prior tax capitalization studies. First, it suggests that institutional investor ownership could mitigate tax capitalization due to firms with greater institutional investor ownership being more liquid. Second, it suggests that prior studies' finding that expected rates of return fell more for non-dividend paying than for dividend-paying firms following JGTRRA03 could be due to non-dividend-paying firms being less liquid.

Key Words: Taxes, Liquidity, Asset Pricing, Expected Return

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1 Introduction

For decades researchers in economics, finance, and accounting have debated and sought evidence that investor-level taxes are impounded into the value of equity shares. We contribute to this debate by offering evidence that liquidity affects the positive relation between investor-level tax rates and expected pretax rates of return that is predicted in many prior studies (e.g., Dhaliwal, Li, Moser, and Krull 2005; Dhaliwal, Krull, and Li 2007; and Scholes, Wolfson, Erickson, Maydew and Shevlin 2009). Our evidence arises from two sources: a theory-based analysis, and empirical findings that are consistent with the predictions of our analysis. Both sources suggest that lower liquidity amplifies and higher liquidity attenuates the general positive relation between investor-level tax rates and firms' expected pretax rates of return. Throughout the remainder of the paper, references to "expected rates of return" are synonymous with references to "expected pretax rates of return."

Although the prediction that tax capitalization is amplified when liquidity is lower and attenuated when liquidity is higher may not be surprising to some readers, to our knowledge no one has posited or tested this prediction. We suspect that this is the result of prior theory-based studies of tax capitalization having been motivated by an after-tax Capital Asset Pricing Model (CAPM) (e.g., Brennan 1970; Gordon and Bradford 1980; Guenther and Sansing 2010). The CAPM is a model of perfect competition; as such, it harbors no notion of liquidity (see, e.g., Lambert, Leuz, and Verrecchia 2011). In contrast, our analysis is motivated by imperfect competition, and liquidity plays a central role in imperfectly competitive markets. The basic intuition that falls out of our analysis is that lower liquidity will amplify the effect of tax rate changes on firms' expected rates of return, whereas higher liquidity will attenuate the effect. Thus, when assessing the behavior of expected rates of return, the effect of a change in a tax rate cannot be divorced from the level of liquidity.

Our analysis relies on several assumptions that facilitate the (purely) mathematical development of our argument. For example, we consider an economy with a single firm whose pretax cash flows are distributed to investors who hold shares in the firm at the end of one period. We also consider a generic, investor-level tax on these pretax cash flows. One could characterize the generic tax as a tax on capital gains or a tax on dividend income. In the empirical tests of our predictions, we utilize changes in both capital gains and dividend tax rates. Assuming that all investors are subject to tax contrasts with recent studies on tax capitalization that allow for some proportion of investors in the economy to be subject to the tax while the remainder are tax-exempt (Guenther and Sansing 2006, 2010; Sikes and Verrecchia 2011). These studies hold that tax capitalization is influenced by the weighted-average tax rate of all investors in the economy (as opposed to the tax status of some marginal investor). Our assumption that all, as opposed to some proportion of, investors are subject to tax has no effect on the results of our analysis. This is because our results concern the *marginal* effect of changes in liquidity and the tax rate on expected rates of return. Changing the mix or proportion of investors in the economy who are subject to tax will affect the magnitude of the impact of a tax rate change on shares prices, but it will not affect the direction of the impact (i.e., positive, negative, or zero).

We center the empirical test of our prediction about the role of liquidity around the

Jobs and Growth Tax Relief Reconciliation Act of 2003 (JGTRRA03), which decreased the maximum statutory individual-level dividend tax rate from 38.1 percent to 15 percent, and the maximum statutory individual-level capital gains tax rate from 20 percent to 15 percent. Prior studies (Dhaliwal, Krull, and Li 2007; Auerbach and Hassett 2007) find that expected rates of return (i.e., cost of capital) fell for both dividend-paying and non-dividend-paying firms following JGTRRA03.[1]

The sample period for our tests is 1999-2007, with 2003 excluded since JGTRRA03 was enacted in May 2003. We use Amihud's (2002) measure of price impact as our measure of illiquidity. Amihud's ratio gives the absolute (percentage) price change per dollar of daily trading volume, or the daily price impact of the order flow. In this sense, the measure is consistent with Kyle's (1985) concept of illiquidity (λ), or the response of price to order flow. Amihud's (2002) measure varies inversely with a firm's level of liquidity. We proxy for firms' expected pretax rates of return with their annual raw buy-and-hold returns. We find that on average returns are significantly higher in the four years following JGTRRA03 than in the four years preceding it. However, our focus is on cross-sectional differences in returns rather than time-series trends in returns. Consistent with the cross-sectional prediction generated from our theoretical analysis, we find that the increase in returns following JGTRRA03 was significantly smaller among less liquid firms. This result holds for both dividend-paying and non-dividend-paying firms.

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[1] Dhaliwal, Krull and Li (2007) compare quarterly implied cost of capital estimates for the six quarters before to the estimates for the six quarters after JGTRRA03 (where they include the enactment quarter in the post period). Auerbach and Hassett (2007) conduct an event study in which they analyze the market reaction over eight five-day windows within which significant news concerning the likelihood of passage of the dividend tax rate cut was made public.

In addition to presenting a new cross-sectional prediction related to tax capitalization, our paper has two implications for prior work on tax capitalization. First, this paper sheds light on the debate regarding whether the degree of dividend tax capitalization is influenced by the tax status of the marginal investor. Several prior studies hold that dividend tax capitalization is stronger the more likely it is that the marginal investor is a taxable investor. These studies use institutional investor ownership to proxy for tax-exempt and corporate ownership and predict that tax capitalization is weaker the greater a firm's institutional investor ownership (e.g., Ayers, Cloyd, and Robinson 2002; Dhaliwal, Li and Trezevant 2003; Dhaliwal, Li, Moser, and Krull 2005; Dhaliwal, Krull, and Li 2007; Campbell, Chyz, Dhaliwal, and Schwartz 2011).[2] Our theoretical and empirical analyses suggest that the mitigating force of institutional investor ownership on dividend tax capitalization documented in prior studies could be the result of an omitted correlated variable: liquidity. Consistent with prior studies, we find that institutional investor ownership mitigates the effect of the 2003 tax rate cuts on expected rates of return. However, once we control for the effect of liquidity on the relation between investor-level tax rates and expected rates of return, we no longer find that institutional ownership plays a mitigating role among dividend-paying firms.

Second, our results have a potentially important implication for prior studies that conclude that expected rates of return decreased more for non-dividend-paying firms than for dividend-paying firms following JGTRRA03. Some view this result in prior studies as surprising based on the expectation that expected rates of return of dividend-paying firms would

[2] Miller and Scholes (1982) also hold that the equilibrium price of the firm is determined by a single marginal investor whose identity determines the pricing of taxes.

be affected by both tax rate cuts, whereas expected rates of return of non-dividend-paying firms would only be affected by the capital gains tax rate cut. One possible explanation for the surprising result is that non-dividend-paying firms are significantly less liquid than dividend-paying firms.

Over the past several decades, research on the economics of dividend taxation has centered on three basic theories: the "tax irrelevance view," the "traditional view," and the "new view." Under the tax irrelevance view, taxable individuals are infra-marginal. In other words, a non-taxable entity (e.g., a pension fund) or symmetrically taxed investor (e.g., a securities dealer) is the relevant price-setter (Black and Scholes 1974; Miller and Scholes 1978, 1982). As a result, changes in the dividend tax rate do not affect expected rates of return, and thereby do not affect investment, payout, and financing decisions. Under the traditional view, there are non-tax benefits from paying dividends and the manager sets dividend policy at the point where the marginal benefit of an extra dollar of dividends equals the marginal tax cost. Reductions in the dividend tax rate can lower the required pretax rate of return. Moreover, because the traditional view assumes that the marginal source of funds for investment is new equity issues (assuming no debt), reductions in the dividend tax rate can also lead to greater investment and higher dividend payouts. The new view assumes that retained earnings are the marginal source of investment funds and all retentions will eventually be distributed as taxable dividends. The market value of equity capitalizes all expected taxes on current and future dividends, even for non-dividend paying firms. The new view predicts that an increase in the dividend tax rate leads to lower equity prices, but dividend yields per se do not explain firm value (because all taxes are already impounded into price, even for non-dividend paying firms) and thus investment and payout decisions are unaffected (Auerbach 1979; Bradford 1981; King 1977). The new view is consistent with what we might expect to occur when a dividend tax rate change is permanent. In reality, the effect of dividend tax rate changes on expected rates of return is likely a mixture of the three different views. On its face, our theoretical analysis is most consistent with the new view in that it does not distinguish between dividend-paying and non-dividend paying firms. It also does not distinguish between tax rates applied to capital gain income versus dividend income, nor does it distinguish between current versus future dividends since we consider an economy with a single firm whose pretax cash flows are distributed at the end of one period to investors who hold shares in the firm. However, our empirical results are also consistent with the traditional view in that, consistent with prior studies, we find that expected rates of return decrease with dividend yield following JGTRRA03.[3]

The paper proceeds as follows. We present our theoretical analysis and its predictions in Section 2. In Section 3, we empirically test the predictions that result from our theoretical analysis. Section 4 concludes.

[3] Hanlon and Heitzman (2010) provide a discussion of the three different views. This paragraph relies heavily on Section 5.1.1 of their paper. Moreover, the number of prior studies on tax capitalization are too numerous for us to review in this paper. We refer readers to Section 5 of Hanlon and Heitzman (2010) for a review of the literature.

2 The effect of liquidity on tax rate changes

Our goal is to understand how liquidity affects the relation between investor-level tax rates and expected pretax rates of return. Toward that goal, first we describe a market setting where liquidity and tax rates play salient roles, and then derive the price of a firm's shares in this setting. Next we study how liquidity affects the impact of tax rate changes on expected rates of return. The results of our analysis suggest that lower liquidity will amplify the impact of tax rate changes, whereas higher liquidity will attenuate the impact of tax rate changes. The theory we provide applies equally to either investor-level dividend tax rates or investor-level capital gains tax rates. Thus, in our analysis below, we characterize "tax" as a generic payment to some governing agency. In Section 3, using changes in the maximum statutory tax rates applied to dividend income and capital gains, we empirically test our prediction that lower liquidity amplifies the effect of tax rate changes on expected rates of return.

2.1 Share price in an imperfectly competitive economy

We consider a one-period economy that consists of one firm, a risk-free asset, and some number of investors. We normalize the (after-tax) return on the risk-free asset to zero and normalize its price at the beginning of the period to 1. Let $S > 0$ represent the number of shares of stock the firm supplies to the economy. Each share of stock generates uncertain cash flow of $\tilde{V}$ at the end of the period, where $\tilde{V} = V$ represents the (per-share) realization of the cash flow. This implies that the firm generates *total* (uncertain) cash flow of $S \cdot \tilde{V}$.

In our analysis the role of S is not benign because as S increases investors in the economy have to bear increasingly more risk. As we show below, this depresses the price of the firm's shares despite the fact that it has no effect on per-share cash flow (i.e., no effect on $\tilde{V}$).

Let P represent the per-share price of the firm's shares at the beginning of the period. We assume there are $N \geq 3$ investors in the economy, each of whom has identical information about the distribution of end of period cash flow.[4] Because investors have homogeneous beliefs, we suppress the specifics of their information and simply assume they believe that the firm's cash flow, $\tilde{V}$, has a normal distribution with expected value $E\left[\tilde{V}\right]$ and variance $Var\left[\tilde{V}\right]$. We assume that each investor has a negative exponential (or CARA) utility function for an amount w given by $-\exp\left[-w/r\right]$, where r is the an investor's constant absolute risk tolerance.

We assume that all investors are subject to tax, and pay tax on the firm's end-of-period cash flow at a rate t. In other words, from the perspective of an individual investor who pays tax, the firm's expected after-tax cash flow is $(1-t)\,E\left[\tilde{V}\right]$ and the variance an investor associates with this after-tax cash flow is $(1-t)^2\,Var\left[\tilde{V}\right]$. As it relates to our assumptions about the role of tax in the economy, we emphasize two points. First, our assumption that all investors are subject to tax, as opposed to some proportion of investors being subject to tax (as in, e.g., Guenther and Sansing 2006, 2010; Sikes and Verrecchia 2011), has no effect on the results we report below. The reason for this is that our results concern the *marginal*

[4] Imperfect competition settings of the type we study require the participation of at least three investors so as to eliminate the possibility of one investor, or a pair of investors, having too much monopoly power: see the discussion on p. 329 of Kyle (1989).

effect of changes in liquidity and the tax rate on expected rates of return. Changing the mix or proportion of investors who are subject to tax will affect the *magnitude* of a tax rate change on share prices, but it plays no role in assessing the direction of the marginal effect (i.e., whether the marginal effect is positive, negative, or zero). In other words, increasing the proportion of investors who are not subject to tax will attenuate the impact of a tax rate change on share prices, but it will not affect the direction of the comparative static that is associated with that change. Second, in a one period economy a firm liquidates at the end of the period and distributes its cash flow proceeds to investors. Hence, in such a setting there is no clear distinction between a tax on dividends versus a tax on capital gains. Consequently, in our analysis we simply characterize the tax as a generic payment to some governing agency paid at the rate of tV, where $\tilde{V} = V$ represents the (per-share) realization of the firm's end-of-period cash flow.

Before proceeding with our analysis, we attempt to describe in broad terms the market process that leads to a determination of firm share price, P. As in Kyle (1989), our analysis characterizes a market process where at the beginning of the period each investor submits a demand function for the firm's shares to a Walrasian auctioneer. The auctioneer aggregates together investors' demands, and then determines a single price for the shares such that at that price demand for shares equals the supply of those shares. As is standard in any rational expectations setting, each investor's demand for shares maximizes an investor's expected utility as a function of the price that clears the market. In this sense our market-clearing process is identical to the vast literature on rational expectations equilibria in market settings.[5] What distinguishes our market-clearing process from the rational expectations literature based on perfect competition, however, is that we assume that each investor's demand function incorporates a belief as to how his demand affects prices, and this belief is self-sustaining in equilibrium.

To characterize the market-clearing process in greater detail, we start by letting D represent an investor's demand for the firm's shares. When investors have a negative exponential utility function and uncertainty is expressed as arising from a normal distribution, the certainty equivalent of each investor's expected utility simplifies into the familiar expression of the expected value of his end-of-period wealth minus a term that is proportional to the variance of his wealth. This implies that each investor chooses D to maximize the following objective function

$$\max_{D} E[(1-t)\tilde{V} - P]D - \frac{1}{2r}D^2(1-t)^2 Var\left[\tilde{V}\right]. \tag{1}$$

Because investors are identically informed (and identical in every other aspect), D is identical across investors; thus there is no need to subscript it further so as to distinguish an individual investor's demand.

In contrast to a model of perfect competition, a model of imperfect competition posits that investors take into consideration the effect of their demand on price. Specifically, and following Kyle (1989) and Lambert et al. (2011), we characterize imperfect competition as the self-sustaining belief by each investor that he faces an upwardly-sloping price curve for

[5] See, for example, Grossman and Stiglitz (1980), Hellwig (1980), Diamond and Verrecchia (1981), Admati (1985), etc.

firm shares. In particular, we assume each investor believes that his demand is related to price as follows:

$$P = p_0 + \lambda \cdot D, \tag{2}$$

where p_0 is the intercept in the determination of P that is unrelated to an investor's demand and λ is the coefficient applied to an investor's demand. We interpret λ as the degree of *illiquidity* associated with an individual investor's demand. When λ is small, an investor's demand moves price less, and thus the market for firm shares is more liquid with respect to demand; when λ is large, an investor's demand moves price more, and thus the market is less liquid for firm shares. Each investor treats λ as fixed when he determines the demand that maximizes his expected utility; our goal below is to determine a self-sustaining λ.

Returning to an investor's optimization problem, an investor solves for D based on knowledge of the realization of the firm share price. In other words, we assume that an investor conditions his expectations, and hence his optimization problem, over P. Conditioning on the market-clearing price implies that an investor who knows P also knows the realization of the intercept p_0 because he chooses D, conjectures λ, and believes that $P = p_0 + \lambda D$. Based on his belief as to how his demand affects prices, an investor solves for D in eqn. (1) by substituting into eqn. (1) the relation $P = p_0 + \lambda D$, taking the derivative of eqn. (1) with respect to D, and then setting the result equal to 0. In other words, an investor takes the derivative of

$$\left((1-t)\,E\left[\tilde{V}\right] - (p_0 + \lambda D)\right) D - \frac{1}{2r} D^2 \,(1-t)^2\, Var\left[\tilde{V}\right]$$

with respect to D and sets the result equal to 0: this yields the first-order condition

$$\left((1-t)\,E\left[\tilde{V}\right]-(p_0+\lambda D)\right)-\lambda D-\frac{1}{r}D\,(1-t)^2\,Var\left[\tilde{V}\right]=0. \tag{3}$$

Substituting P for the relation $p_0+\lambda D$ back into the expression above and solving for the investor's optimal demand yields

$$D=\left(\frac{1}{r}\,(1-t)^2\,Var\left[\tilde{V}\right]+\lambda\right)^{-1}\left((1-t)\,E\left[\tilde{V}\right]-P\right). \tag{4}$$

Market clearing requires that investors' demand for firm shares equals the supply of those shares. That is, P must satisfy

$$N\cdot D\,(P)-S=0,$$

where here we characterize an investor's demand, $D\,(\cdot)$, as a function of the price of the firm's shares, P. Substituting for $D\,(P)$ from eqn. (4) yields the following result:

$$P=(1-t)\,E\left[\tilde{V}\right]-\left(\frac{1}{r}\,(1-t)^2\,Var\left[\tilde{V}\right]+\lambda\right)\frac{S}{N}. \tag{5}$$

We codify our analysis to this point in the following proposition.

Proposition 1. *In an imperfectly competitive economy where N identically informed investors compete to hold shares in a firm, the price of the firm's shares is*

$$P=(1-t)\,E\left[\tilde{V}\right]-\left(\frac{1}{r}\,(1-t)^2\,Var\left[\tilde{V}\right]+\lambda\right)\frac{S}{N}.$$

Proposition 1 implies that the firm's share price can be expressed as three components. The first component, $(1-t)\,E\left[\tilde{V}\right]$, is expected after-tax cash flow. The second component, $\frac{1}{r}\,(1-t)^2\,Var\left[\tilde{V}\right]\frac{S}{N}$, is the discount in price that manifests in a setting where competition among risk averse investors to hold firm shares is perfect. The third component, $\lambda \frac{S}{N}$, is an *additional* discount in the expression for price that measures the extent to which the market is illiquid (i.e., the extent to which the market is not perfectly competitive). As we alluded to above, both discounts increase as S increases because investors in the economy have to bear increasingly more risk.

To expand on this issue briefly, consider the possibility that the economy is perfectly competitive. Perfect competition is tantamount to assuming $\lambda = 0$, and here the price of the firm's shares in eqn. (5) reduces to

$$P = (1-t)\, E\left[\tilde{V}\right] - \left(\frac{1}{r}(1-t)^2\, Var\left[\tilde{V}\right]\right)\frac{S}{N}.$$

Alternatively, if competition is imperfect then this is tantamount to assuming $\lambda > 0$ and thus price absorbs the additional discount $\lambda \frac{S}{N}$: see, for example, the discussion in Lambert et al. (2011).

2.2 Solving for λ

The next step in the analysis is to solve for λ. Although the solution follows directly from Kyle (1989) (see also Lambert et al. 2011) and thus is straightforward, nonetheless it is complicated in detail. Thus, we relegate the solution to the Appendix and here simply state the result: solving for λ yields

$$\lambda = \frac{1}{N-2}\frac{1}{r}(1-t)^2\, Var\left[\tilde{V}\right]. \tag{6}$$

To digress briefly, when $\lambda = \frac{1}{N-2}\frac{1}{r}(1-t)^2 Var\left[\tilde{V}\right]$ one can also show using the results in the Appendix that in equilibrium each investor holds

$$D = \frac{S}{N}$$

shares in the firm. This result is intuitive: in equilibrium, each of N identical investors holds $1/N^{-th}$ of the total number of the firm's shares, S. This implies that each investor bears the risk associated with uncertain cash flow of $\frac{S}{N} \cdot \tilde{V}$, and in aggregate investors bear the risk of the firm's total cash flow of $S \cdot \tilde{V}$.

We codify our analysis to this point with the following result.

Proposition 2. *In an imperfectly competitive economy where N identically informed investors compete to hold shares in a firm, one can show that*

$$\lambda = \frac{1}{r(N-2)}(1-t)^2 Var\left[\tilde{V}\right],$$

and thus the price of the firm's shares reduces to

$$P = (1-t)E\left[\tilde{V}\right] - \frac{N-1}{rN(N-2)}(1-t)^2 Var\left[\tilde{V}\right] \cdot S.$$

Note that the expression for price in Proposition 2 is unique and well-defined, provided that $N \geq 3$.

Our next step is to use the results of this section to study expected pretax rates of return.

2.3 Comparative statics

Using the expression for the price of the firm's shares in Proposition 2, the firm's expected pretax rate of return is

$$
\begin{aligned}
E\left[\frac{\tilde{V}-P}{P}\right] &= \frac{E\left[\tilde{V}\right]-P}{P} \\
&= \frac{tE\left[\tilde{V}\right]+\frac{(N-1)(1-t)^2 Var\left[\tilde{V}\right]}{rN(N-2)}S}{(1-t)\,E\left[\tilde{V}\right]-\frac{(N-1)(1-t)^2 Var\left[\tilde{V}\right]}{rN(N-2)}S}. \qquad (7)
\end{aligned}
$$

Our goal is to understand how liquidity affects the impact of a change in the tax rate on the firm's expected pretax rate of return. Toward achieving that goal, first we have to address three issues.

The first issue is that λ in our analysis, which is sometimes referred to as "Kyle's λ" in reference to Kyle (1985, 1989), is a measure of *illiquidity*. Consequently, we re-state our goal as one of determining the marginal effect of a change in *illiquidity* on the marginal effect of a change in the tax rate on the firm's expected pretax rate of return. This goal can be expressed mathematically as one of determining the sign of the cross-partial derivative

$$\frac{\partial}{\partial \lambda}\frac{\partial}{\partial t}E\left[\frac{\tilde{V}-P}{P}\right].$$

To digress briefly, if the cross-partial derivative is continuous, as is the case in our analysis, then the order in which we take a derivative makes no difference: in other words,

$$\frac{\partial}{\partial \lambda}\frac{\partial}{\partial t}E\left[\frac{\tilde{V}-P}{P}\right]=\frac{\partial}{\partial t}\frac{\partial}{\partial \lambda}E\left[\frac{\tilde{V}-P}{P}\right].$$

Thus, an alternative interpretation of our analysis below is that we study the marginal effect of a change in the tax rate on the marginal effect of a change in illiquidity on the firm's expected pretax rate of return.

The second issue is that in our analysis λ is an endogenous variable, and so properly we should not be using it to take a derivative. A technique for addressing this issue is to reverse the roles of λ and N. Specifically, N is ostensibly an exogenous parameter, but by appealing to eqn. (6) we can define N in terms of λ, as opposed to defining λ in terms of N: in other words, eqn. (6) implies

$$N = \frac{(1-t)^2 \, Var\left[\tilde{V}\right]}{r\lambda} + 2. \tag{8}$$

This technique treats N as an endogenous variable and λ as an exogenous parameter. In turn, using eqns. (7) and (8), we can re-express the firm's expected pretax return in terms of λ as

$$E\left[\frac{\tilde{V}-P}{P}\right] = \frac{tE\left[\tilde{V}\right] + \frac{\lambda\left(r\lambda+(1-t)^2 Var\left[\tilde{V}\right]\right)}{2r\lambda+(1-t)^2 Var\left[\tilde{V}\right]} S}{(1-t)\,E\left[\tilde{V}\right] - \frac{\lambda\left(r\lambda+(1-t)^2 Var\left[\tilde{V}\right]\right)}{2r\lambda+(1-t)^2 Var\left[\tilde{V}\right]} S}. \tag{9}$$

This allows us to compute $\frac{\partial}{\partial\lambda}\frac{\partial}{\partial t}E\left[\frac{\tilde{V}-P}{P}\right]$ by representing $E\left[\frac{\tilde{V}-P}{P}\right]$ by the right-hand-side of eqn. (9).

To digress briefly, note that if we require that $N \geq 3$ for the expression for price in Proposition 2 to be unique and well defined, then in eqn. (8) it must be the case that

$$\frac{(1-t)^2 \, Var\left[\tilde{V}\right]}{r\lambda} \geq 1,$$

or

$$\lambda \leq \frac{(1-t)^2 \, Var\left[\tilde{V}\right]}{r}. \tag{10}$$

Thus, one feature to the approach of treating λ as an exogenous parameter is that it restricts λ to levels of illiquidity that are not "excessively large," where the latter is defined by the right-hand-side of eqn. (10). That said, this restriction is simply the reciprocal of the requirement that at least three investors compete to hold firm shares (i.e., $N \geq 3$), which, arguably, is not an especially burdensome, real-world, institutional assumption.

The third issue is that we would like to eliminate from consideration potentially perverse cases that arise from the possibility that the price of the firm's shares is negative. We interpret a negative share price as a circumstance where investors have to be *subsidized* to hold a firm's shares - once again, an unlikely outcome in real-world institutional settings. A necessary condition that share price be nonnegative is that end-of-period expected cash flow is nonnegative: that is, $E\left[\tilde{V}\right] \geq 0$. Even in the presence of nonnegative expected cash flow, however, share price can be negative if, for example: 1) the firm's expected cash flow is too low in relation to the variance of its cash flow (i.e., $E\left[\tilde{V}\right]$ is low and $Var\left[\tilde{V}\right]$ high); 2) investors' tolerance for risk is too low (i.e., r is low); and/or 3) the supply of the firm's shares is too high (i.e., S is high). Recall that we assume that $S > 0$. Thus, an inequality that restricts the supply of the firm's shares in relation to other (exogenous) parameters so as to ensure that share price is nonnegative is

$$0 < S \leq \frac{(1-t)\,E\left[\tilde{V}\right]}{\frac{\lambda\left(r\lambda+(1-t)^2 Var\left[\tilde{V}\right]\right)}{2r\lambda+(1-t)^2 Var\left[\tilde{V}\right]}}. \tag{11}$$

Henceforth we assume this to be the case. For convenience we let S_0 represent the right-hand-side of eqn. (11):

$$S_0 \equiv \frac{(1-t)\,E\left[\tilde{V}\right]}{\frac{\lambda\left(r\lambda+(1-t)^2 Var\left[\tilde{V}\right]\right)}{2r\lambda+(1-t)^2 Var\left[\tilde{V}\right]}}.$$

This allows us to restate eqn. (11) as the requirement that $0 < S \leq S_0$.

Having addressed these three issues, next we calculate $\frac{\partial}{\partial\lambda}\frac{\partial}{\partial t}E\left[\frac{\tilde{V}-P}{P}\right]$ under the assumption firm share price is nonnegative (i.e., $S \leq S_0$). The calculation is computationally very tedious and so we sketch its derivation. First determine $\frac{\partial}{\partial\lambda}\frac{\partial}{\partial t}E\left[\frac{\tilde{V}-P}{P}\right]$ and then observe that it can be decomposed into the product of two mathematical expressions. The first expression is nonnegative provided that $S \leq S_0$. The second expression is monotonically decreasing in S. Let $F(S)$ represent the second expression as a function of S. Because $F(S)$ is monotonically decreasing in S, $F(S) \geq F(S_0)$ and we show

$$\begin{aligned} F(S) &\geq F(S_0) \\ &\propto \left(2r^2\lambda^2 + r\lambda(1-t)^2 Var\left[\tilde{V}\right] + \left((1-t)^2 Var\left[\tilde{V}\right]\right)^2\right) \\ &\quad \times \left(2r^2\lambda^2 + 2r\lambda(1-t)^2 Var\left[\tilde{V}\right] + \left((1-t)^2 Var\left[\tilde{V}\right]\right)^2\right) \\ &> 0, \end{aligned}$$

where the symbol "$\propto$" denotes proportionality. This establishes that we can express $\frac{\partial}{\partial\lambda}\frac{\partial}{\partial t}E\left[\frac{\tilde{V}-P}{P}\right]$ as the product of two terms, both of which are nonnegative, and hence $\frac{\partial}{\partial\lambda}\frac{\partial}{\partial t}E\left[\frac{\tilde{V}-P}{P}\right]$ is nonnegative. Consequently, the marginal effect of a change in illiquidity on the marginal effect of a change in the tax rate on expected pretax rates of return is nonnegative (provided that share price is nonnegative). We state this formally as our third proposition.

Proposition 3. *In an imperfectly competitive economy where N identically informed (and* *endogenously determined) investors compete to hold shares in a firm and the resulting price of those shares is nonnegative, the marginal effect of a change in illiquidity on the marginal effect of a change in the tax rate on the firm's expected pretax rate of return is nonnegative.*

The economic intuition that underlies Proposition 3 is that increased illiquidity amplifies the effect of an increase in the tax rate on expected pretax rates of return. This implies that in assessing the behavior of expected pretax rates of return, the effect of a change in a tax rate cannot be divorced from the level of liquidity. Lower liquidity will amplify the effect of tax rate changes, whereas higher liquidity will attenuate the effect of tax rate changes.

3 Empirical Tests

In this section, we empirically test our prediction that lower liquidity amplifies and higher liquidity attenuates the positive relation between investor-level tax rates and expected pretax rates of return. Our tests are centered around the Jobs and Growth Tax Relief Reconciliation Act of 2003 (JGTRRA03), which decreased the maximum statutory individual-level dividend tax rate from 38.1 percent to 15 percent and the maximum statutory individual-level long-term capital gains tax rate from 20 percent to 15 percent. Prior studies (Dhaliwal, Krull, and Li 2007; Auerbach and Hassett 2007) find that expected pretax rates of return (i.e., cost of capital) fell for both dividend-paying and non-dividend-paying firms. We begin by estimating the following Ordinary Least Squares (OLS) regression to test whether expected rates of return declined more for less liquid firms following JGTRRA03:

$$
\begin{aligned}
&RETURN_{i,y} = \beta_1 + \beta_2 POSTACT_y + \beta_3 ILLIQUIDITY_{i,y-1} + \\
&\beta_4 ILLIQUIDITY_{i,y-1} * POSTACT_y + \beta_5 INST_{i,y-1} + \\
&\beta_6 SIZE_{i,y-1} + \beta_7 BM_{i,y-1} + \beta_8 DISPERSION_{i,y-1} + \beta_9 ROE_{i,y-1} + \\
&\beta_{10} LEV_{i,y-1} + \beta_{11} BETA\_MKTRF_{i,y-1} + \beta_{12} BETA\_SMB_{i,y-1} + \\
&\beta_{13} BETA\_HML_{i,y-1} + \beta_{14} BETA\_UMD_{i,y-1} + \beta_{15-44} INDUSTRY_{i,y} + \varepsilon. \quad (12)
\end{aligned}
$$

A negative β_4 will be consistent with our prediction.

Our proxy for expected rates of return is firm i's buy-and-hold return (including dividends) calculated over year y ($RETURN$). The sample period for the tests is 1999-2007, with 2003 excluded since the tax rate cuts were enacted in May 2003. $POSTACT$ equals one for years 2004-2007 and equals zero for years 1999-2002. We estimate equation (12) first including all firms, then separately for dividend-paying firms only and for non-dividend-paying firms only. In equation (12) and all other regressions in the paper, we cluster standard errors by firm.

Our proxy for lower levels of liquidity is Amihud's (2002) measure of price impact. It equals the ratio of the daily absolute return to the dollar trading volume on that day, averaged over the trading days in year *y-1* for which there is return and volume data, or

$$ILLIQUIDITY_{i,y-1} = \frac{1}{D_{i,y-1}} \sum_{d=1}^{D_{i,y-1}} \left(1,000 * \sqrt{\frac{|R_{i,y-1,d}|}{VOLD_{i,y-1,d}}} \right)$$

where $|R_{i,y\text{-}1,d}|$ is firm i's absolute return on day d of year *y-1*, $VOLD_{i,y-1,d}$ is the respective daily volume in dollars, and $D_{i,y-1}$ is the number of days for which data are available for firm i in year *y-1.* Amihud's ratio gives the absolute (percentage) price change per dollar of daily trading volume, or the daily price impact of the order flow. In this sense, the measure is consistent with Kyle's (1985) concept of illiquidity (λ), or the response of price to order flow. Similar to Chen, Goldstein, and Jiang (2010), we use the square root version of the Amihud (2002) measure. Consistent with Amihud (2002), we apply several restrictions when calculating the ratio. First, we require for there to be stock return and volume data for more than 200 days of year *y-1* for firm i in order to calculate its ratio. Second, we require for the stock price to be greater than \$5 at the end of year y-1. Third, data to calculate firm i's market capitalization must be available at the end of year y-1. Fourth, after applying the restrictions above, we winsorize observations for the measure at the 1st and 99th percentiles for each year.

Consistent with prior literature, we control for the impact of information asymmetry on a firm's expected rate of return. We include several controls for information asymmetry. The first is analyst forecast dispersion, measured as the standard deviation of analysts' forecasts of firm i's one-year-ahead earnings per share divided by the absolute value of the mean forecast of firm i's one-year-ahead earnings per share (with zero mean forecast observations excluded from the sample) (Diether, Malloy, and Scherbina 2002; Johnson 2004; Sadka and Scherbina 2007). The measure is right-skewed; thus, we use the natural logarithm of the measure (*DISPERSION*). Other control variables that prior studies find are potentially associated with information asymmetry are the percent of outstanding shares owned by institutional investors (*INST*), firm size measured as the natural logarithm of market capitalization (*SIZE*), book-to-market ratio (*BM*), and leverage measured as the sum of current and long-term liabilities scaled by total assets (*LEV*). One might expect for information asymmetry to be lower among larger firms and firms with greater institutional investor ownership and higher among firms with higher book-to-market ratios and higher leverage (Sadka and Scherbina 2007). We control for profitability with the ratio of net income before extraordinary items divided by market capitalization (*ROE*). We control for risk by including the coefficient estimates from estimating a four-factor Fama-French-Carhart model (Fama and French 1993; Carhart 1997) using return data from the 48 months prior to year *y-1* (*BETA_ MKTRF*, *BETA_ SMB*, *BETA_ HML*, and *BETA_ UMD*). We control for industry effects by including an indicator variable for each of Fama and French's 30 industry portfolios (Fama and French 1997). With the exception of the *POSTACT* indicator variable and the industry indicator variables, we measure all independent variables in year y-1. We winsorize all continuous variables used in our analysis at the 1st and 99th percentiles. Table 1 includes variable definitions.

[INSERT TABLE 1 HERE]

Our sample includes firms that are listed on either the New York Stock Exchange (NYSE), American Stocks Exchange (AMEX), or NASDAQ; are in the Center for Research in Security Prices (CRSP)/Compustat Merged Database; have data on quarterly institutional holdings reported on Form 13F available in the Thomson-Reuters database; and are followed by three for more analysts and have analyst forecast data in IBES. Table 2 reports the descriptive statistics for the variables in equation (12). All values are reported in decimal format.

[INSERT TABLE 2 HERE]

Columns (1), (2), and (3) of Table 3 present the results of estimating equation (12) for all firms, dividend-paying firms only, and non-dividend-paying firms only, respectively. The positive coefficient on *POSTACT* in all three columns is due to the fact that realized returns are significantly higher in the four years following than in the four years preceding JGTRRA03. Untabulated tests show that the mean *RETURN* equals 0.017 over the years 1999-2002 and equals 0.078 over the years 2004-2007. Our prediction relates to cross-sectional differences, not to time-series trends in returns. In all three columns, consistent with our prediction, we find that the coefficient on the interaction *ILLIQUIDITY*POSTACT* is negative and significant at the 1% level. Although returns were on average higher in the four years following than in the four years preceding JGTRRA03, this result suggests that the increase in returns was significantly smaller for less liquid firms.

In addition to being statistically significant, the results are economically meaningful. When the analysis is conducted using all firms, the results suggest that a one standard deviation increase (decrease) in illiquidity results in a 4.1 percentage point larger decrease (increase) in annual returns in the four years following 2003 relative to the four years preceding 2003. [6] When the analysis is conducted using dividend-paying firms only, the results suggest that a one standard deviation increase (decrease) in illiquidity results in a 3.2 percentage point larger decrease (increase) in annual returns in the four years following 2003 relative to the four years preceding 2003. When the analysis is conducted using non-dividend-paying

[6] We calculate the 4.1 percentage point change as the coefficient estimate on *ILLIQUIDITY*POSTACT* in column (1) (-0.3293) multiplied by standard deviation of *ILLIQUIDITY* (0.124).

firms only, the results suggest that a one standard deviation increase (decrease) in illiquidity results in a 4.6 percentage point larger decrease (increase) in annual expected returns in the four years following 2003 relative to the four years preceding 2003.[7]

The coefficient on *ILLIQUIDITY* is positive and significant in all three columns, consistent with expected rates of return being positively related to illiquidity (see, e.g., Amihud and Mendelson 1986; Brennan and Subrahmanyam 1996; Brennan, Chordia, and Subrahmanyam 1998; Datar, Naik, and Radcliffe 1998; Haugen and Baker 1996, and Hu 1997 with respect to firm-level liquidity, and Pastor and Stambaugh 2003 with respect to market-wide liquidity). With the exception of the insignificant coefficient on *SIZE* in column (2), consistent with expected rates of return being lower among larger firms and firms with greater institutional ownership, the coefficients on *SIZE* and *INST* are negative and significant. Consistent with information asymmetry being greater among firms with higher book-to-market ratios and with greater analyst forecast dispersion, the coefficient on *BM* is positive and significant in columns (1)-(2) and the coefficient on *DISPERSION* is positive and significant in column (2).

[INSERT TABLE 3 HERE]

We next examine 1) whether expected rates of return decrease with dividend yield following JGTRRA03, and 2) if so, whether the decrease is greater among less liquid firms. To

[7] The calculation is analogous to the one for all firms. Among the dividend-paying (non-dividend-paying) firms in our sample, the standard deviation of *ILLIQUIDITY* equals 0.104 (0.135).

test 1), we estimate equation (13) below:

$$RETURN_{i,y} = \beta_1 + \beta_2 POSTACT_y + \beta_3 YIELD_{i,y-1} +$$
$$\beta_4 YIELD_{i,y-1} * POSTACT_y + \beta_5 HIGHILLIQUIDITY_{i,y-1} + \beta_6 INST_{i,y-1} +$$
$$\beta_7 SIZE_{i,y-1} + \beta_8 BM_{i,y-1} + \beta_9 DISPERSION_{i,y-1} + \beta_{10} ROE_{i,y-1} +$$
$$\beta_{11} LEV_{i,y-1} + \beta_{12} BETA\_MKTRF_{i,y-1} + \beta_{13} BETA\_SMB_{i,y-1} +$$
$$\beta_{14} BETA\_HML_{i,y-1} + \beta_{15} BETA\_UMD_{i,y-1} + \beta_{16-45} INDUSTRY_{i,y} + \varepsilon. \quad (13)$$

The variable *YIELD* equals the amount of dividends paid to common shareholders in year $y-1$ scaled by firm i's market capitalization, both collected from Compustat (Ayers, Cloyd, and Robinson 2002). The variable *HIGHILLIQUIDITY* equals one if firm i's value of *ILLIQUIDITY* is greater than the median value across the sample for the year, and equal to zero otherwise. All other variables are previously defined. We estimate equation (13) first including all firms and setting *YIELD* equal to zero for non-dividend-paying firms and then only including dividend-paying firms. The mean (median) *YIELD* equals 0.01 (0.00) for all firms 0.02 (0.02) for dividend-paying firms only (untabulated). We demean the value of *YIELD* used in the interaction term by subtracting the sample mean.[8] Columns (1) and (2) of Table 4 report the results of estimating equation (13) for all firms and for dividend-paying firms only, respectively. The coefficient on the interaction *YIELD*POSTACT* is negative and significant at the 1% level, consistent with expected rates of return decreasing more following JGTRRA03 for higher dividend yield firms.

[8] Demeaning, or centering, continuous variables used in interactions reduces the amount of multicollinearity that is induced by multiplying together two independent variables. See Aiken and West (1991) for a discussion of the benefits of demeaning variables.

We next estimate the following regression to test whether the change in the relation between dividend yield and expected rates of return following JGTRRA03 is greater among less liquid firms:

$$RETURN_{i,y} = \beta_1 + \beta_2 POSTACT_y + \beta_3 YIELD_{i,y-1} + \beta_4 YIELD_{i,y-1} * POSTACT_y +$$
$$\beta_5 HIGHILLIQUIDITY_{i,y-1} + \beta_6 HIGHILLIQUIDITY_{i,y-1} * POSTACT_y +$$
$$\beta_7 YIELD_{i,y-1} * POSTACT_y * HIGHILLIQUIDITY_{i,y-1} +$$
$$\beta_8 INST_{i,y-1} + \beta_9 SIZE_{i,y-1} + \beta_{10} BM_{i,y-1} + \beta_{11} DISPERSION_{i,y-1} +$$
$$\beta_{12} ROE_{i,y-1} + \beta_{13} LEV_{i,y-1} + \beta_{14} BETA\_MKTRF_{i,y-1} + \beta_{15} BETA\_SMB_{i,y-1} +$$
$$\beta_{16} BETA\_HML_{i,y-1} + \beta_{17} BETA\_UMD_{i,y-1} + \beta_{18-47} INDUSTRY_{i,y} + \varepsilon. \tag{14}$$

We demean the value of *YIELD* used in both interaction terms by subtracting the sample mean. Columns (3) and (4) of Table 4 report the results of estimating equation (14) for all firms and for dividend-paying firms only, respectively. Consistent with the change in the relation between expected rates of return and dividend yield following JGTRRA03 being greater for less liquid firms, the coefficient on the interaction *YIELD*POSTACT*HIGHILLIQUIDITY* is negative; however, it is not statistically significant (p-value=0.29 in column (3) and 0.23 in column (4)).

[INSERT TABLE 4 HERE]

In summary, the results in Table 3 are consistent with the predictions generated from our theoretical analysis in Section 2. Although returns are higher in the four years following than in the four years preceding JGTRRA03, the increase in returns is significantly smaller among less liquid firms. In addition to presenting a new cross-sectional prediction related to tax capitalization, our paper has two important implications for prior work on tax capitalization. We address each of the implications below.

3.1 Marginal Investor

Several prior studies predict that the dividend tax capitalization effect is stronger the more likely it is that the marginal investor is a taxable investor (e.g., Dhaliwal, Li and Trezevant 2003; Ayers, Cloyd, and Robinson 2002; Dhaliwal, Krull, Li and Moser 2005; Dhaliwal, Krull, and Li 2007; Campbell, Chyz, Dhaliwal, and Schwartz 2011). These studies use the percent of a firm's outstanding shares owned by institutional investors to proxy for tax-exempt and corporate ownership and predict that institutional investor ownership mitigates dividend tax capitalization (i.e., dividend tax capitalization is weaker the more likely it is that a tax-exempt or corporate investor is the marginal investor setting price). Such a prediction is inconsistent with prior theoretical studies that rely upon the after-tax CAPM. The after-tax CAPM predicts that the weighted average tax rate of all investors in the economy (where the weight depends on investors' risk tolerances) rather than the tax rate of the marginal investor in the firm is the relevant tax rate in determining the extent of tax capitalization (see Guenther and Sansing (2010) and Hanlon and Heitzman (2010) for a discussion of the issue).

Our theoretical analysis and empirical findings suggest that prior studies' finding that institutional investor ownership mitigates dividend tax capitalization could be due to an omitted correlated variable problem: stocks with greater institutional investor ownership are generally more liquid. Over our sample period, we find that the correlation between *INST* and *ILLIQUIDITY* is negative and statistically significant (untabulated Pearson correlation $= -0.45$, $p < 0.001$). To investigate the issue further, we next estimate a modified version of equation (12) where we first interact *POSTACT* with *INST* rather than with *ILLIQUIDITY*, and then include both interactions, *INST*POSTACT* and *ILLIQUIDITY*POSTACT*. We demean the values of *INST* and *ILLIQUIDITY* used in the interaction terms by subtracting their sample means. We estimate both specifications twice: once including all firms and once only including dividend-paying firms. A positive coefficient on *INST*POSTACT* will be consistent with the finding in prior studies that institutional ownership mitigates dividend tax capitalization. However, if the finding in prior studies that institutional ownership mitigates dividend tax capitalization is due to an omitted correlated variable (i.e., liquidity), then when we include the interaction *ILLIQUIDITY*POSTACT*, the coefficient on *INST*POSTACT* should no longer be positive.

Columns (1)-(2) and (3)-(4) of Table 5 report the results for all firms and for dividend-paying firms only, respectively. In columns (1)-(2), the coefficient on *INST*POSTACT* is positive and significant at the 1% level, regardless of whether the interaction *ILLIQUIDITY*POSTACT* is included in the estimation. Thus, the results for all firms do not suggest that prior studies suffer from an omitted correlated variable. It is possible that dividend tax capitalization is stronger among dividend-paying firms than among non-dividend-paying firms. Therefore, we next just analyze dividend-paying firms. In column (3), the coefficient on *INST*POSTACT* is positive and significant at the 1% level, consistent with findings in prior studies. However, the interaction is no longer significant in column (4) when we also include the interaction *ILLIQUIDITY*POSTACT*. The results when we only include dividend-paying firms suggest that prior studies that conclude that institutional ownership mitigates dividend tax capitalization due to institutional investors being tax-insensitive suffer from an omitted correlated variable. Once one controls for the effect of liquidity on the relation between tax rates and expected rates of return, institutional ownership no longer appears to play a mitigating role, at least among dividend-paying firms. In both columns (2) and (4), the coefficient on *ILLIQUIDITY*POSTACT* is negative and significant at the 5% level, consistent with our prediction.

[INSERT TABLE 5 HERE]

In summary, our results for dividend-paying firms in Table 5 suggest that prior studies that attribute the mitigating force of institutional ownership on dividend tax capitalization to the marginal investor being tax-insensitive suffer from an omitted correlated variable problem. Once we control for the effect of liquidity on the relation between investor-level tax rates and expected rates of return, we no longer find that institutional ownership mitigates dividend tax capitalization. We are not the first to disagree with the marginal investor interpretation. Prior theoretical studies posit that the dividend tax capitalization effect is a function of the weighted-average tax rate across *all* investors, where the weight depends on investors' tolerances for risk (e.g., Brennan 1970; Gordon and Bradford 1980; Michaely and Vila 1995; Guenther and Sansing 2010; Bond, Devereux, and Klemm 2007). According to Guenther and Sansing (2010), one reason that prior empirical studies find that institutional holdings mitigate the dividend tax capitalization effect is that institutional holdings are negatively correlated with individual (i.e., taxable) investors' tolerances for risk. Guenther and Sansing's (2010) suggestion that differences in investors' risk tolerances explain the relation between institutional holdings and dividend tax capitalization and our suggestion that liquidity explains this relation are not mutually exclusive explanations.

3.2 Dividend vs. Non-Dividend Paying Firms

Auerbach and Hassett (2007) and Dhaliwal, Krull, and Li (2007) document that expected rates of return decrease more for non-dividend-paying firms than for dividend-paying firms following JGTRRA03. Some view this result as surprising based on the expectation that expected rates of return of dividend-paying firms would be affected by both tax rate cuts, whereas expected rates of return of non-dividend-paying firms would only be affected by the capital gains tax rate cut. Auerbach and Hassett (2007) and Chen, Dai, Shackelford and Zhang (2011) offer possible explanations for the unexpected result. Auerbach and Hassett (2007) explain that if investors viewed the dividend tax rate cut as permanent (or at least semi-permanent), then investors could have incorporated expected future dividend taxes when pricing the shares of firms that currently were not paying a dividend but which investors expected to pay a dividend prior to a potential reversal in the dividend tax rate cut. Chen, Dai, Shackelford, and Zhang (2011) provide empirical evidence suggesting that expected rates of return fell more for non-dividend-paying firms than for dividend-paying firms because the former are more financially constrained and thus have the most pressing need for capital.

We offer a third possible explanation for the result. Consistent with the prediction from our theoretical analysis, we posit that expected rates of return could have decreased more for non-dividend-paying than for dividend-paying firms following JGTRRA03 due to the fact that non-dividend-paying firms are less liquid. In other words, prior studies that conclude that expected rates of return decreased more for non-dividend-paying firms following JGTRRA03 could suffer from an omitted correlated variable problem. In untabulated tests, we find that the non-dividend-paying firms are significantly less liquid than the dividend-paying firms in our sample. The mean (median) $ILLIQUIDITY$ equals 0.12 (0.08) for the non-dividend paying firms and 0.08 (0.04) for the dividend-paying firms, and the differences between the means and medians are statistically significant at the one percent level.

We first examine whether, like Auerbach and Hassett (2007) and Dhaliwal, Krull, and Li (2007), we find that expected rates of return fall more for non-dividend-paying firms than for dividend-paying firms following JGTRRA03. We estimate the following OLS regression:

$$\begin{aligned}
&RETURN_{i,y} = \beta_1 + \beta_2 POSTACT_y + \beta_3 DIV_{i,y-1} \\
&+\beta_4 DIV_{i,y-1} * POSTACT_y + \beta_5 ILLIQUIDITY_{i,y-1} + \beta_6 INST_{i,y-1} + \\
&\beta_7 SIZE_{i,y-1} + \beta_8 BM_{i,y-1} + \beta_9 DISPERSION_{i,y-1} + \beta_{10} ROE_{i,y-1} + \\
&\beta_{11} LEV_{i,y-1} + \beta_{12} BETA_MKTRF_{i,y-1} + \beta_{13} BETA_SMB_{i,y-1} + \\
&\beta_{14} BETA_HML_{i,y-1} + \beta_{15} BETA_UMD_{i,y-1} + \beta_{16-45} INDUSTRY_{i,y} + \varepsilon. \quad (15)
\end{aligned}$$

The variable DIV is an indicator variable that equals one if firm i pays a dividend to common shareholders in year y-1; zero otherwise. All other variables are previously defined. A positive β_4 will be consistent with the finding in Auerbach and Hassett (2007) and Dhaliwal, Krull, and Li (2007). After estimating equation (15) as outlined above, we re-estimate it with the inclusion of the interaction *ILLIQUIDITY*POSTACT*. If the finding in prior studies that expected rates of return decreased more for non-dividend-paying firms than for dividend-paying firms following JGTRRA03 is due to an omitted correlated variable, then once we include *ILLIQUIDITY*POSTACT* in the estimation, the coefficient on *DIV*POSTACT* should no longer be positive.

Column (1) of Table 6 presents the results of estimating equation (15) without the inclusion of the interaction *ILLIQUIDITY*POSTACT*. The coefficient on *DIV*POSTACT* is insignificant. Thus, inconsistent with prior studies, we do not find a significant difference in the change in expected rates of return following JGTRRA03 between dividend-paying and non-dividend-paying firms. The difference in results could be due to different samples and/or methodologies. For example, Dhaliwal, Krull and Li (2007) compare quarterly implied cost of capital estimates for the six quarters before to the estimates for the six quarters after JGTRRA03 (where they include the enactment quarter in the post period). Auerbach and Hassett (2007) conduct an event study in which they analyze the market reaction over eight five-day windows within which significant news concerning the likelihood of passage of the dividend tax rate cut was made public.

In column (2) we add the interaction *ILLIQUIDITY*POSTACT*. Consistent with our prediction, the coefficient on *ILLIQUIDITY*POSTACT* is negative and significant at the 1% level. The coefficient on *DIV*POSTACT* is now positive and significant at the 10% level. This result suggests that once one controls for the impact of liquidity, expected rates of return appear to have decreased more for dividend-paying firms than for non-dividend-paying firms following JGTRRA03. However, we place little emphasis on this result since the coefficient on *DIV*POSTACT* is only marginally significant. Moreover, due to our inability in column (1) to replicate the result in prior studies that expected rates of return declined more for non-dividend-paying firms following JGTRRA03, we cannot conclude that controlling for liquidity would alter the results in prior studies.

In summary, the results in this section suggest that the puzzling result in prior studies that expected rates of return decreased more for non-dividend-paying firms than for dividend-paying firms following JGTRRA03 could be due to prior studies not controlling for differences in liquidity among dividend-paying and non-dividend-paying firms. We will continue to investigate this issue.

[INSERT TABLE 6 HERE]

4 Conclusion

This paper offers a new theory related to cross-sectional variation in tax capitalization. Unlike prior theory studies on dividend tax capitalization that rely on the after-tax CAPM, our analysis features an imperfectly competitive market in which liquidity plays a role in determining the effect of investor-level tax rates on firms' expected pretax rates of return. The prediction generated from our analysis is that lower liquidity amplifies the effect of a change in an investor-level tax rate on expected rates of return, whereas higher liquidity attenuates the effect. Our theory-based analysis includes a generic investor-level tax rate, which can be generalized to both investor-level dividend and capital gains tax rates.

We empirically test our prediction using the cuts to the maximum statutory tax rate on dividend income (from 38.1 to 15 percent) and on capital gain income (from 20 percent to 15 percent), enacted by the Jobs and Growth Tax Relief Reconciliation Act of 2003. We find that annual buy-and-hold returns are significantly higher in the four years following than in the four years preceding JGTRRA03. Consistent with our cross-sectional prediction, we find that the increase in returns following JGTRRA03 is significantly smaller among less liquid firms.

This paper results in a couple of important implications for prior studies. First, this paper contributes to the recent debate regarding whether the identity of the marginal investor determines the extent to which investor-level taxes are priced. Our results suggest that prior studies that attribute the attenuating force of institutional ownership on dividend tax capitalization to the marginal investor being tax-exempt (or tax-insensitive with respect to dividends) could suffer from an omitted correlated variable problem. Once we control for the effect of liquidity on the relation between investor-level tax rates and expected rates of return, we no longer find that institutional ownership attenuates dividend tax capitalization among dividend-paying firms. Second, prior studies conclude that expected rates of return decreased more for non-dividend-paying firms than for dividend-paying firms following JGTRRA03. Some view this result as puzzling based on the expectation that expected rates of return of dividend-paying firms would be affected by both tax rate cuts, whereas expected rates of return of non-dividend-paying firms would only be affected by the capital gains tax rate cut. We posit a possible explanation for the puzzling result. Consistent with the prediction from our theoretical analysis, perhaps expected rates of return decreased more for non-dividend-paying than for dividend-paying firms following JGTRRA03 due to the fact that non-dividend-paying firms are less liquid.

Appendix

The solution for λ

To proceed, first we provide an alternative characterization of the investor's demand strategy. In particular, in competing with other informed investors, we conjecture that each investor adopts a strategy

$$D = \alpha - \boldsymbol{\beta} \cdot P,$$

where α is an intercept term and β is the weight an investor places on P. For this conjectured strategy to be rational based on the computation of D in eqn. (4), it must be the case that

$$\boldsymbol{\beta} = \left(\frac{1}{r} (1-t)^2 \, Var \left[\tilde{V} \right] + \lambda \right)^{-1} \text{ and}$$

$$\alpha = \left(\frac{1}{r} (1-t)^2 \, Var \left[\tilde{V} \right] + \lambda \right)^{-1} (1-t) \, E \left[\tilde{V} \right].$$

Market clearing requires that investors' demand for firm shares equals the supply of those shares. Thus, P must satisfy

$$N \cdot D(P) - S = 0,$$

where once again $D(P)$ represents an investor's demand as a function of P. Substituting $D = \alpha - \beta P$ into this equation and solving for the P that clears the market yields

$$P = \Delta (N\alpha - S),$$

where Δ is given by

$$\Delta = \frac{1}{N} \beta^{-1} = \frac{1}{N} \left(\frac{1}{r} (1-t)^2 \, Var \left[\tilde{V} \right] + \lambda \right).$$

Here, Δ represents the marginal impact on price of an additional share brought to (or withdrawn from) the market by the firm.

To ensure the market clearing condition $P = \Delta(N\alpha - S)$ in conjunction with the conjecture that $D = \alpha - \beta P$, p_0 and λ in the expression $P = p_0 + \lambda D$ must be of the form

$$p_0 = \lambda((N-1)\alpha - S) \text{ and } \lambda = \frac{1}{N-1}\beta^{-1},$$

respectively. The rationale for this claim is that when this is the case

$$\begin{aligned} P &= p_0 + \lambda D \\ &= \lambda((N-1)\alpha - S) + \lambda(\alpha - \beta P) \\ &= \lambda(N\alpha - S) - \lambda\beta P, \end{aligned}$$

which implies $P(1 + \lambda\beta) = \lambda(N\alpha - S)$ and thus

$$\frac{\lambda}{1 + \lambda\beta} = \Delta = \frac{1}{N}\beta^{-1}; \tag{A1}$$

solving for λ in eqn. (A1) yields $\lambda = \frac{1}{N-1}\beta^{-1}$. We further solve for λ by substituting in for β in the identity

$$\lambda = \frac{1}{N-1}\beta^{-1} = \frac{1}{N-1}\left(\frac{1}{r}(1-t)^2 Var\left[\tilde{V}\right] + \lambda\right).$$

This yields the result:

$$\lambda = \frac{1}{N-2}\frac{1}{r}(1-t)^2 Var\left[\tilde{V}\right].$$

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Table 1
Variable Definitions

RETURN	=	firm *i*'s annual raw buy-and-hold return (including dividends)
POSTACT	=	1 for years 2004-2007 and 0 for years 1999-2002 (2003 is excluded since the tax rate cuts were enacted in May 2003)
ILLIQUIDITY	=	square root version of Amihud's (2002) measure of price impact. See text for calculation.
YIELD	=	firm *i*'s dividends paid to common shareholders (DVC from Compustat) divided by firm *i*'s market capitalization (CSHO*PRCC_F from Compustat)
DIV	=	1 if *YIELD* > 0; 0 otherwise
INST	=	% of firm *i*'s outstanding years owned by institutional investors
SIZE	=	natural log of firm *i*'s market capitalization
BM	=	firm *i*'s book equity (CEQ from Compustat) divided by market capitalization (CSHO*PRCC_F from Compustat)
DISPERSION	=	natural logarithm of the standard deviation of analysts' forecasts of firm *i*'s one-year-ahead earnings per share divided by the absolute value of the mean forecast of one-year-ahead earnings per share (with zero mean forecast observations excluded from the sample)
ROE	=	firm *i*'s net income before extraordinary items (IBCOM from Compustat) divided by market capitalization (CSHO*PRCC_F from Compustat)
LEV	=	sum of firm *i*'s current and long-term liabilities (DLC+DLTT from Compustat) scaled by total assets (AT from Compustat)
BETA_MKTRF	=	firm *i*'s loading on MKTRF from estimating Fama-French-Carhart 4-factor model using return data for the prior 48 months
BETA_SMB	=	firm *i*'s loading on SMB from estimating Fama-French-Carhart 4-factor model using return data for the prior 48 months
BETA_HML	=	firm *i*'s loading on HML from estimating Fama-French-Carhart 4-factor model using return data for the prior 48 months
BETA_UMD	=	firm *i*'s loading on UMD from estimating Fama-French-Carhart 4-factor model using return data for the prior 48 months

Table 2
Descriptive Statistics

Variable	Mean	Std. Dev.	5th Pctile	25th Pctile	Median	75th Pctile	95th Pctile
RETURN	0.049	0.563	-0.673	-0.250	0.004	0.243	0.869
POSTACT	0.514	0.500	0.000	0.000	1.000	1.000	1.000
ILLIQUIDITY	0.103	0.124	0.010	0.028	0.060	0.130	0.346
INST	0.607	0.237	0.177	0.435	0.633	0.792	0.965
SIZE	20.848	1.538	18.557	19.727	20.673	21.801	23.761
BM	0.469	0.334	0.078	0.252	0.417	0.613	1.044
DISPERSION	-3.584	1.239	-5.242	-4.485	-3.801	-2.890	-1.200
ROE	0.019	0.142	-0.143	0.012	0.043	0.066	0.115
LEV	0.230	0.207	0.000	0.040	0.200	0.355	0.620
BETA_MKTRF	1.075	0.969	-0.093	0.574	1.002	1.512	2.583
BETA_SMB	0.684	1.173	-0.640	0.053	0.514	1.162	2.612
BETA_HML	0.221	1.421	-2.158	-0.330	0.386	0.962	2.025
BETA_UMD	-0.143	0.862	-1.449	-0.479	-0.092	0.216	0.984

This table presents the descriptive statistics for variables used in equation (12) (N=17,011). The sample period is 1999-2007, excluding 2003. See Table 1 for variable definitions. All values are in decimal format.

Table 3
Effect of Liquidity on the Relation between Tax Rates and Expected Rates of Return

Variables	(1) All Firms	(2) Dividend-Paying Firms	(3) Non-Dividend Paying Firms
POSTACT	0.0759***	0.0644***	0.0957***
	(9.008)	(7.019)	(6.895)
ILLIQUIDITY	0.1971***	0.1540**	0.1614*
	(2.775)	(2.040)	(1.655)
*ILLIQUIDITY*POSTACT*	-0.3293***	-0.3045***	-0.3433***
	(-3.695)	(-3.311)	(-2.703)
INST	-0.0578**	-0.0554***	-0.0652*
	(-2.480)	(-2.686)	(-1.672)
SIZE	-0.0113***	-0.0019	-0.0310***
	(-2.755)	(-0.478)	(-4.170)
BM	0.0379**	0.0605***	0.0241
	(2.186)	(3.202)	(0.983)
DISPERSION	0.0055	0.0082*	0.0049
	(1.270)	(1.795)	(0.751)
ROE	0.1877***	0.3745***	0.1598**
	(3.327)	(5.561)	(2.429)
LEV	0.0009	-0.0708***	0.0523
	(0.039)	(-2.967)	(1.390)
BETA_MKTRF	-0.0152***	-0.0008	-0.0185***
	(-2.835)	(-0.092)	(-2.966)
BETA_SMB	-0.0184***	-0.0296***	-0.0151**
	(-3.177)	(-3.826)	(-2.222)
BETA_HML	0.0178***	0.0324***	0.0132**
	(3.687)	(4.427)	(2.296)
BETA_UMD	-0.0168**	-0.0315***	-0.0117
	(-2.354)	(-2.826)	(-1.417)
Constant	0.2516***	0.0589	0.6327***
	(2.633)	(0.611)	(3.760)
Observations	17,011	7,863	9,148
R-squared	0.023	0.056	0.021

Columns (1), (2), and (3) present the results of estimating equation (12) for all firms, dividend-paying firms only, and non-dividend paying firms only, respectively. The sample period is 1999-2007, with 2003 excluded. See Table 1 for variable definitions. Industry indicator variables are included in the estimation but suppressed in the table. T-statistics are calculated using standard errors clustered by firm and appear in parentheses below coefficient estimates. ***,**, * denote statistical significant at $p < 0.01$, $p < 0.05$, and $p < 0.10$, respectively, using a two-tailed test.

Table 4
Effect of Liquidity on the Relation between Tax Rates, Dividend Yield, and Expected Rates of Return

Variables	(1) All Firms	(2) Dividend-Paying Firms	(3) All Firms	(4) Dividend-Paying Firms
POSTACT	0.0857***	0.0928***	0.1329***	0.1218***
	(10.502)	(8.655)	(13.152)	(10.630)
YIELD	0.5775*	1.2112***	0.7620**	1.1647***
	(1.930)	(3.262)	(2.536)	(3.104)
*YIELD*POSTACT*	-1.3395***	-1.6678***	-1.4320***	-1.2604***
	(-3.845)	(-3.908)	(-3.727)	(-2.701)
HIGHILLIQUIDITY	-0.0648***	-0.0050	-0.0137	0.0449***
	(-4.828)	(-0.418)	(-0.746)	(2.798)
*HIGHILLIQUIDITY*POSTACT*			-0.0973***	-0.0848***
			(-5.794)	(-4.572)
*YIELD*POSTACT*HIGHILLIQUIDITY*			-0.4438	-0.6504
			(-1.059)	(-1.206)
INST	-0.0914***	-0.0516**	-0.0881***	-0.0488**
	(-3.865)	(-2.468)	(-3.734)	(-2.361)
SIZE	-0.0289***	-0.0037	-0.0288***	-0.0041
	(-6.550)	(-0.983)	(-6.534)	(-1.080)
BM	0.0419**	0.0563***	0.0343**	0.0461**
	(2.439)	(2.933)	(1.985)	(2.393)
DISPERSION	0.0064	0.0089*	0.0067	0.0084*
	(1.478)	(1.943)	(1.539)	(1.834)
ROE	0.1999***	0.3759***	0.1920***	0.3654***
	(3.559)	(5.454)	(3.401)	(5.433)
LEV	0.0091	-0.0753***	0.0050	-0.0760***
	(0.352)	(-2.873)	(0.197)	(-2.929)
BETA_MKTRF	-0.0171***	0.0029	-0.0170***	0.0027
	(-3.154)	(0.314)	(-3.155)	(0.300)
BETA_SMB	-0.0175***	-0.0278***	-0.0177***	-0.0272***
	(-3.003)	(-3.577)	(-3.031)	(-3.487)
BETA_HML	0.0183***	0.0319***	0.0186***	0.0313***
	(3.757)	(4.364)	(3.828)	(4.301)
BETA_UMD	-0.0147**	-0.0311***	-0.0156**	-0.0320***
	(-2.046)	(-2.793)	(-2.172)	(-2.880)
Constant	0.6863***	0.0855	0.6584***	0.0742
	(6.473)	(0.946)	(6.202)	(0.828)
Observations	17,011	7,863	17,011	7,863
R-squared	0.024	0.057	0.026	0.061

Columns (1)-(2) and (3)-(4) present the results of estimating equations (13) and (14), respectively. The sample includes all firms in columns (1) and (3) and dividend-paying firms only in columns (2) and (4). The sample period is 1999-2007, with 2003 excluded. See Table 1 for variable definitions. Industry indicator variables are included in the estimation but suppressed in the table. T-statistics are calculated using standard errors clustered by firm and appear in parentheses below coefficient estimates. ***,**, * denote statistical significant at $p < 0.01$, $p < 0.05$, and $p < 0.10$, respectively, using a two-tailed test.

Table 5
Effect of Institutional Ownership on the Relation between Tax Rates and Expected Rates of Return

Variables	(1) All Firms	(2) All Firms	(3) Dividend-Paying Firms	(4) Dividend-Paying Firms
POSTACT	0.0841***	0.0809***	0.0785***	0.0679***
	(10.012)	(9.231)	(8.845)	(6.607)
INST	-0.1462***	-0.1268***	-0.1037***	-0.0793**
	(-3.749)	(-3.184)	(-3.474)	(-2.558)
*INST*POSTACT*	0.1753***	0.1302***	0.0901***	0.0408
	(4.344)	(3.068)	(2.666)	(1.046)
ILLIQUIDITY	0.0960	0.1570**	0.0568	0.1363*
	(1.539)	(2.149)	(0.828)	(1.745)
*ILLIQUIDITY*POSTACT*		-0.2187**		-0.2638**
		(-2.342)		(-2.523)
SIZE	-0.0095**	-0.0108***	-0.0002	-0.0018
	(-2.362)	(-2.632)	(-0.049)	(-0.451)
BM	0.0425**	0.0393**	0.0675***	0.0615***
	(2.474)	(2.262)	(3.636)	(3.240)
DISPERSION	0.0045	0.0049	0.0076*	0.0079*
	(1.044)	(1.143)	(1.662)	(1.729)
ROE	0.1949***	0.1910***	0.3766***	0.3740***
	(3.462)	(3.388)	(5.504)	(5.543)
LEV	0.0051	0.0019	-0.0672***	-0.0697***
	(0.206)	(0.077)	(-2.804)	(-2.916)
BETA_MKTRF	-0.0156***	-0.0155***	-0.0006	-0.0011
	(-2.913)	(-2.906)	(-0.061)	(-0.125)
BETA_SMB	-0.0180***	-0.0185***	-0.0286***	-0.0295***
	(-3.126)	(-3.209)	(-3.672)	(-3.807)
BETA_HML	0.0179***	0.0182***	0.0319***	0.0322***
	(3.687)	(3.757)	(4.341)	(4.400)
BETA_UMD	-0.0162**	-0.0168**	-0.0314***	-0.0318***
	(-2.264)	(-2.339)	(-2.817)	(-2.852)
Constant	0.2693***	0.2791***	0.0531	0.0700
	(2.797)	(2.882)	(0.552)	(0.726)
Observations	17,011	17,011	7,863	7,863
R-squared	0.023	0.024	0.055	0.056

Columns (1)-(2) and (3)-(4) report the results of estimating modified versions of equation (12) including all firms and dividend-paying firms only, respectively. The sample period is 1999-2007, with 2003 excluded. See Table 1 for variable definitions. Industry indicator variables are included in the estimation but suppressed in the table. T-statistics are calculated using standard errors clustered by firm and appear in parentheses below coefficient estimates. ***,**, * denote statistical significant at $p < 0.01$, $p < 0.05$, and $p < 0.10$, respectively, using a two-tailed test.

Table 6
Relation between Tax Rates and Expected Rates of Return:
Dividend-Paying vs. Non-Dividend-Paying Firms

Variables	(1)	(2)
POSTACT	0.0853***	0.0887***
	(6.453)	(6.721)
DIV	0.0144	0.0217*
	(1.120)	(1.677)
*DIV*POSTACT*	-0.0141	-0.0275*
	(-0.936)	(-1.827)
ILLIQUIDITY	0.1124*	0.2054***
	(1.803)	(2.864)
*ILLIQUIDITY*POSTACT*		-0.3505***
		(-3.918)
INST	-0.0490**	-0.0590**
	(-2.112)	(-2.508)
SIZE	-0.0099**	-0.0119***
	(-2.416)	(-2.848)
BM	0.0422**	0.0369**
	(2.454)	(2.126)
DISPERSION	0.0052	0.0056
	(1.199)	(1.296)
ROE	0.1913***	0.1856***
	(3.385)	(3.283)
LEV	0.0062	0.0009
	(0.252)	(0.036)
BETA_MKTRF	-0.0147***	-0.0148***
	(-2.711)	(-2.748)
BETA_SMB	-0.0169***	-0.0179***
	(-2.877)	(-3.048)
BETA_HML	0.0165***	0.0171***
	(3.365)	(3.499)
BETA_UMD	-0.0158**	-0.0166**
	(-2.203)	(-2.314)
Constant	0.2196**	0.2522***
	-2.301	(2.604)
Observations	17,011	17,011
R-squared	0.022	0.023

This table reports the results of estimating equation (15) including all firms. The sample period is 1999-2007, with 2003 excluded. See Table 1 for variable definitions. Industry indicator variables are included in the estimation but suppressed in the table. T-statistics are calculated using standard errors clustered by firm and appear in parentheses below coefficient estimates.***,**, * denote statistical significant at $p < 0.01$, $p < 0.05$, and $p < 0.10$, respectively, using a two-tailed test.