



## The on-the-run liquidity phenomenon<sup>☆</sup>

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### ABSTRACT

We test the implications of a model of multi-asset speculative trading in which liquidity differentials between on-the-run and off-the-run U.S. Treasury bonds ensue from endowment shocks in the presence of two realistic market frictions—information heterogeneity and imperfect competition among informed traders—and a public signal. Our evidence suggests that (i) off/on-the-run liquidity differentials are economically and statistically significant, even after controlling for several of the bonds' intrinsic characteristics (such as duration, convexity, repo rates, or term premiums), and (ii) off/on-the-run liquidity differentials are smaller immediately following bond auction dates, and larger when the uncertainty surrounding the ensuing auction allocations is high, when the dispersion of beliefs across informed traders is high, and when macroeconomic announcements are noisy, consistent with our model.

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### 1. Introduction

The *on-the-run phenomenon* refers to the stylized fact that, in fixed income markets, securities with nearly identical cash flows trade at different yields and with different liquidity. In particular, most recently issued (i.e., *on-the-run*, *new*, or *benchmark*) government bonds of a certain maturity are generally more expensive and liquid

than previously issued (i.e., *off-the-run* or *old*) bonds maturing on similar dates.

Ample evidence of this phenomenon has been reported both in the U.S. Treasury market (e.g., Amihud and Mendelson, 1991; Kamara, 1994; Furfine and Remolona, 2002; Krishnamurthy, 2002; Strebulaev, 2002; Fleming, 2003; Goldreich, Hanke and Nath, 2005) and in other countries (e.g., for Japan, Mason, 1987; Boudouck and Whitelaw, 1991, 1993). Accordingly, several explanations have also been provided by practitioners and academics. The most popular one attributes off/on-the-run yield differentials to liquidity—the extent to which an asset can be traded cheaply, quickly, and with limited price impact. The liquidity premium hypothesis of Amihud and Mendelson (1986) states that since investors value liquidity, more liquid securities should trade at a premium over otherwise similar, yet less liquid ones. Most existing literature concentrates on testing this prediction. Early studies find support for it (e.g., Amihud and Mendelson, 1991; Warga, 1992; Kamara, 1994). More recent research suggests that off/on-the-run yield differentials may be explained by such considerations as differing tax

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treatments (Strebulaev, 2002), specialness in the repo markets (i.e., the cost of shorting, as in Duffie, 1996; Krishnamurthy, 2002), search costs (Vayanos and Weill, 2008), or the value of future liquidity (Goldreich, Hanke and Nath, 2005).

In spite of this debate on the extent of off/on-the-run yield differentials and the relative importance of liquidity as an explanatory factor (the on-the-run *price* phenomenon), there is little or no disagreement in the literature that off/on-the-run liquidity differentials (the on-the-run *liquidity* phenomenon) are both economically and statistically significant (e.g., Amihud and Mendelson, 1991; Strebulaev, 2002). Nonetheless, we are aware of no theoretical and empirical study of the determinants of those liquidity differentials.<sup>1</sup> Performing such an analysis is the objective of this paper.<sup>2</sup> To that purpose, we develop a parsimonious model of multi-asset trading. The model—in the spirit of Kyle (1985), Foster and Viswanathan (1996), and Pasquariello and Vega (2007)—builds upon two realistic market frictions: information heterogeneity and imperfect competition among informed traders (henceforth, speculators). In this basic setting, speculators trade strategically based on their private signals. This leads uninformed market-makers (MMs) to worsen equilibrium market liquidity. More diverse information among speculators makes their trading activity more cautious and MMs more vulnerable to adverse selection. This leads to even lower equilibrium market liquidity. Pasquariello and Vega (2007) find strong empirical support for these implications of the model in the U.S. Treasury market.<sup>3</sup>

We use this setting to identify a novel mechanism explaining the on-the-run liquidity phenomenon. Specifically, we explore the role of government auctions in discriminating among two asset types of identical terminal payoff, off-the-run and on-the-run bonds, since by definition the latter are those most recently auctioned to sophisticated traders. In addition, the individual allocations these traders receive from the auction process are unknown to market participants. We capture these features of government bond markets by further assuming that each speculator receives an uninformative, privately observed endowment shock in the on-the-run asset and cares about the interim as well as the liquidation value of his portfolio. In this amended setting, we show that

equilibrium market liquidity in the on-the-run asset is greater than in the off-the-run asset, the more so the greater the uncertainty about endowment shocks. Intuitively, speculators trade strategically in the on-the-run asset based not only on their private signals (as in the off-the-run asset) but also on their endowment shocks. The latter ameliorates adverse selection in on-the-run trading and induces the MMs to make the on-the-run market more liquid than the off-the-run market.

As interestingly, the resulting equilibrium off/on-the-run liquidity differential is sensitive to the information environment in which trading takes place. In particular, we show that such differential is generally lower the more correlated speculators' private fundamental information is. More homogeneous private signals attenuate speculators' incentives to trade cautiously in both markets; yet they alleviate adverse selection the most where it is most severe (i.e., in the off-the-run market). Consistently, we also show that, *ceteris paribus*, the equilibrium off/on-the-run liquidity differential is decreased by the availability of public fundamental news—a trade-free source of information about assets' payoffs reducing the adverse selection risk for the MMs—the more so the greater is that signal's precision.

The contribution of the model is twofold. Other papers have studied the properties of a financial market in which strategic traders receive privately observable endowment shocks, most notably Vayanos (1999, 2001), and Bhattacharyya and Nanda (2008). Yet, to our knowledge, our model is the first to relate off/on-the-run liquidity differentials to auction-driven endowment shocks.<sup>4</sup> Furthermore, our model is the first to generate explicit and empirically testable implications on the impact of both the heterogeneity of private signals and the presence and quality of public signals on the nature of that relationship.

Our empirical results strongly support the main implications of our model. We start by providing additional evidence of the on-the-run liquidity phenomenon in the U.S. Treasury market.<sup>5</sup> We show that daily averages of intraday bid–ask spread differentials between the second most recently auctioned (i.e., *just* off-the-run) three-month, six-month, and one-year Treasury bills, and two-year, five-year, and 10-year Treasury notes and the corresponding on-the-run securities are positive, economically significant—averaging more than half of the corresponding mean off-the-run spread—and cannot be explained by differences in such fundamental characteristics of the underlying securities as modified duration, convexity, repo differentials, and term premiums. Our

<sup>1</sup> Amihud and Mendelson (1991) and Vayanos and Weill (2008) report anecdotal evidence that off-the-run bonds are in smaller effective supply, hence less liquid, because they become locked away in institutional investors' portfolios. Barclay, Hendershott and Kotz (2006) show that the market share of electronic trading platforms drops significantly when Treasury securities go off-the-run. Those platforms were not available during most of our sample period.

<sup>2</sup> A related literature studies price discrepancies among substantially identical securities or portfolios (e.g., Lee, Schleifer, and Thaler, 1990, 1991; Daves and Ehrhardt, 1993; Bodurtha, Kim, and Lee, 1995; Froot and Dabora, 1999; Grinblatt and Longstaff, 2000). Many of these papers use liquidity differentials to explain observed mispricings, yet none examines directly the determinants of those differentials.

<sup>3</sup> Consistently, Sadka and Scherbina (2007) find a positive relationship between analyst disagreement and both the permanent price impact of trades and the effective percentage bid–ask spread in the U.S. equity market.

<sup>4</sup> Nyborg and Strebulaev (2004) explore the strategic behavior of bidders with exogenous, pre-auction long/short endowment shocks in multiunit uniform and discriminatory auctions when short squeezing can occur in the secondary market. See also Nyborg and Strebulaev (2001).

<sup>5</sup> Amihud and Mendelson (1991) find that the difference between the relative bid–ask yield spread of U.S. Treasury notes and bills with matched maturities of less than six months is about 2.25%. Strebulaev (2002) finds similarly large absolute bid–ask yield spread differentials when comparing U.S. Treasury notes with different initial maturity but maturing on the same day.

analysis suggests that these off/on-the-run liquidity differentials are affected by uncertainty about speculators' endowments in the on-the-run securities, consistent with our model. In particular, we find that in the days immediately following Treasury "new bond" auction dates—when on-the-run endowment uncertainty is arguably the highest—off/on-the-run bid–ask spread differentials are smaller, often significantly so, even after controlling for relative duration, convexity, repo specialness, and supply effects. Accordingly, we also show that off/on-the-run liquidity differentials are positively related to the competitive yield range (high minus low divided by average auction bid yield), a more direct proxy for auction-driven endowment uncertainty.

Further investigation reveals that the magnitude and dynamics of those liquidity differentials are also crucially related to the informational role of trading in the U.S. Treasury market, again consistent with our model. In particular, we find that off/on-the-run spread differentials are positively related to perceived, marketwide uncertainty surrounding U.S. monetary policy—measured by Eurodollar implied volatility—and to the degree of information heterogeneity about U.S. macroeconomic fundamentals among market participants—measured by the standard deviation of professional forecasts of macroeconomic news releases (as in Pasquariello and Vega, 2007)—albeit more weakly so. Correspondingly, we show that the availability of macroeconomic news lowers off/on-the-run bid–ask spread differentials, the more so when those signals are less noisy and/or when speculators' private information is more heterogeneous.

We proceed as follows. In Section 2, we construct a stylized model of trading to guide our empirical analysis. In Section 3, we describe the data. In Section 4, we present the empirical results. We conclude in Section 5.

## 2. A model of the on-the-run liquidity phenomenon

The objective of our study is to propose and test a novel explanation of the on-the-run liquidity phenomenon in the *secondary* U.S. Treasury bond market—one based on both endowment uncertainty from the *primary* market for government bonds and adverse selection from post-auction trading. The primary market is where the U.S. Treasury sells securities, in "astonishing" quantity (\$3.42 trillion in calendar year 2003, according to Garbade and Ingber, 2005), to the public: retail and institutional investors, specially designated large players, known as *primary dealers*, and any broker or dealer acting on behalf of customers (Fabozzi and Fleming, 2004). The secondary market for Treasury securities is among the largest, most active, and most liquid financial markets. Trading in this market occurs in an interdealer over-the-counter setting in which primary and non-primary dealers act as MMs, trading with customers on their own accounts and among themselves via *interdealer brokers*.<sup>6</sup>

We begin our investigation by developing the simplest stylized representation of the process of price formation in the secondary Treasury bond market apt for our objective. Specifically, we first describe a parsimonious model of trading in on-the-run and off-the-run securities in the spirit of Subrahmanyam (1991), Foster and Viswanathan (1996), and Pasquariello and Vega (2007), and derive closed-form solutions for the equilibrium depth differential between the two assets in the presence of endowment shocks to the former. Then, we enrich the model by introducing a public signal (e.g., macroeconomic news) and consider its implications for the market equilibrium. We test for the statistical and economic significance of our theoretical argument in the remainder of the paper. All proofs are in the Appendix.

### 2.1. The basic model

The basic model is a three-date, two-period economy in which two identical risky assets ( $i = 1, 2$ ) are exchanged. Trading occurs only at the end of the first period ( $t = 1$ ). At the end of the second period ( $t = 2$ ), the identical payoff of the risky assets—a normally distributed random variable  $v$  with mean  $p_0$  and variance  $\sigma_v^2$ —is realized. The economy is populated by three types of risk-neutral traders: a discrete number ( $M$ ) of informed traders (that we label speculators), liquidity traders, and perfectly competitive MMs in each asset  $i$ . All traders know the structure of the economy and the decision process leading to order flow and prices. At time  $t = 0$  there is no information asymmetry about  $v$ , and the price of both risky assets is  $p_0$ .

In fixed income markets, just-issued, on-the-run government bonds (e.g., asset 2) routinely trade at different prices and with different liquidity than previously issued, off-the-run bonds with (almost) identical cash flows (e.g., asset 1). In this section, we propose a theory of the latter phenomenon that focuses on the crucial role of government auctions in discriminating among these assets. Indeed, by definition, on-the-run Treasury bonds are so by having been most recently auctioned to sophisticated traders in the primary market. There is a significant body of research studying the various processes through which multiple identical units such as government securities are sold (e.g., see Krishna, 2002, for a review). During the 1990s, the U.S. Treasury moved from a discriminatory (i.e., multiple-price) to a uniform-price auction format to sell its securities to the public, following numerous violations of auction rules in 1991 (Garbade and Ingber, 2005). The implications of this decision for auction revenues, the likelihood of short squeezes in the post-auction market, post-auction volatility, or pre-auction (*when-issued*) pricing are at the center of a lively debate in the literature (e.g., Back and Zender, 1993; Nyborg and Sundaresan, 1996; Chatterjea and Jarrow, 1998; Nyborg and Strebulaev, 2004; Goldreich, 2007).<sup>7</sup> Nonetheless, in both settings the individual

<sup>6</sup> For more details on the microstructure of the U.S. Treasury market, see Fabozzi and Fleming (2004) and Mizrach and Neely (2007).

<sup>7</sup> The when-issued market is an active forward market in Treasury securities soon to be auctioned (e.g., Mizrach and Neely, 2007).

allocations of Treasury securities resulting from the auction process (or a portion thereof) may be unknown to all other bidders and market participants (e.g., Back and Zender, 1993).

We capture these features of the primary government bond market by assuming that, at time  $t = 0$ , each speculator  $k$  receives an initial endowment of risky asset 2 whose magnitude  $e_2(k)$ —a normally distributed random variable with mean  $\bar{e}$  and variance  $\sigma_e^2$ —is known exclusively to him. Because of this assumption, we label asset 2 the on-the-run security in our setting. We can interpret  $\bar{e}$  as the expected auction outcome for that speculator and  $e_2(k) - \bar{e}$  as his positive or negative, auction-driven exogenous endowment shock in asset 2. To reduce notation, we impose that  $\bar{e} = 0$ . Individual allocation shocks are endogenous in a number of auction models. For instance, Back and Zender (1993) extend auction theory by deriving equilibrium outcomes for risk-neutral, heterogeneously informed bidders in discriminatory and uniform auctions of Treasury securities as perfectly divisible goods with uncertain intrinsic value. More recently, Nyborg and Strebulaev (2004) explore the strategic behavior of auction bidders faced with the possibility of a short squeeze in the post-auction market for an asset of known intrinsic value. In their framework, auction participants receive exogenous, common knowledge long/short endowment shocks from the pre-auction market, but the ensuing auction outcomes are endogenously determined in equilibrium. We concentrate on the impact of exogenous endowment shocks in one asset on the strategic behavior of informed traders in secondary markets for all assets.<sup>8</sup> In that respect, these and other auction models may provide a further rationale for the sign of and uncertainty about those shocks in our setting. We also assume that these endowment shocks are independent ( $cov[e_2(k), e_2(j)] = 0$ ) and uninformative about  $v$  ( $cov[e_2(k), v] = 0$ ), hence so is each speculator's initial wealth  $W_0(k) = e_2(k)p_0$ . Individual auction outcomes are likely to be related to each other or to the unobservable unit value of the asset sold (e.g., Back and Zender, 1993). However, similar yet more involved results ensue if either  $cov[e_2(k), e_2(j)] \neq 0$  or  $cov[e_2(k), v] \neq 0$  (e.g., see Pasquariello, 2003).

Sometime between  $t = 0$  and 1, each speculator  $k$  also receives a private and noisy signal of  $v$ ,  $S_v(k)$ . We assume that each signal  $S_v(k)$  is drawn from a normal distribution with mean  $p_0$  and variance  $\sigma_s^2$  and that, for any two  $S_v(k)$  and  $S_v(j)$ ,  $cov[v, S_v(k)] = cov[S_v(k), S_v(j)] = \sigma_v^2$  and  $cov[e_2(k), S_v(k)] = cov[e_2(k), S_v(j)] = 0$ . The analysis that follows would be similar but more complex if  $cov[v, S_v(k)] \neq cov[S_v(k), S_v(j)]$ , as in Foster and Viswanathan (1996) and Pasquariello and Vega (2007). We can interpret  $S_v(k)$  as private information (or private interpretation of public information) about any factor or state variable determining the future resale value of Treasury securities. For instance, Brandt and Kavajecz (2004, p. 2624) observe that

Treasury market participants' subjective valuations of the traded securities may be due to their own models for the current state of the economy, level and dynamics of the yield curve, or their interaction, as well as to some truly private information (e.g., as in the case of "a hedge fund with an ex-member of the Federal Reserve Board").<sup>9</sup> Brandt and Kavajecz (2004), Green (2004), and Pasquariello and Vega (2007) provide strong evidence for the informational role of trading in the process of price formation in the secondary market for Treasury securities (even in the absence of public information releases). Our assumptions imply that  $E[v|S_v(k)] - p_0 = \delta_v(k) = \rho[S_v(k) - p_0]$ , where  $\rho = \sigma_v^2/\sigma_s^2$  is the correlation between any two information endowments  $\delta_v(k)$  and  $\delta_v(j)$ . We parametrize the degree of diversity among speculators' signals by imposing that  $\sigma_s^2 = \sigma_v^2/\rho$  and  $\rho \in (0, 1]$ . If  $\rho = 1$ , speculators' private information is homogeneous, i.e., all speculators receive the same signal  $S_v(k) = S_v$ . If  $\rho < 1$ , speculators' information is heterogeneous, i.e., less than perfectly correlated, the more so the lower is  $\rho$ . Evidence for significant and persistent differences in information among traders—perhaps due to different sources (e.g., about fundamentals or past trading activity), skills, processing abilities, or resources—is commonly found in most financial markets (e.g., Diether, Malloy, and Scherbina, 2002; Sadka and Scherbina, 2007; Kallberg and Pasquariello, 2008), including those for government bonds (e.g., Pasquariello and Vega, 2007).

### 2.1.1. Market participants and trading

At time  $t = 1$ , both speculators and liquidity traders submit their orders in assets 1 and 2 to the MMs, before these assets' equilibrium prices  $p_{t,i}$  have been set. We define the market order of speculator  $k$  in asset  $i$  to be  $x_i(k)$ . Liquidity traders generate random, normally distributed demands  $z_1$  and  $z_2$ , with mean zero and variance  $\sigma_z^2$ . For simplicity, we assume that  $z_1$  and  $z_2$  are independent from all other random variables. By the same token, we also impose that perfectly competitive MMs in each asset  $i$  do not receive any information about its terminal payoff  $v$ , but observe only that asset's aggregate order flow  $\omega_{1,i} = \sum_{k=1}^M x_i(k) + z_i$  (as in Subrahmanyam, 1991) before setting the market-clearing price  $p_{1,i} = p_{1,i}(\omega_{1,i})$ . This latter assumption allows for the possibility that the equilibrium prices of assets 1 and 2 be different in equilibrium. It can be relaxed to let the MMs observe the aggregate order flow for all securities (i.e.,  $p_{1,1} = p_{1,1}(\omega_{1,1}, \omega_{1,2})$  and  $p_{1,2} = p_{1,2}(\omega_{1,1}, \omega_{1,2})$ , as in Caballé and Krishnan, 1994; Pasquariello, 2003, 2007) if their terminal payoffs are similar yet not identical. The ensuing setting, albeit more complex, yields similar equilibrium implications.<sup>10</sup>

<sup>9</sup> Evans (2002) and Evans and Lyons (2002, 2008), Evans and Lyons (2008) make a similar argument for currency markets. Albuquerque, de Francisco and Marques (2008) show that order flow in the U.S. equity market contains marketwide private information.

<sup>10</sup> Since the MMs in either asset do not possess private information about  $v$  and hold their positions until liquidation (at  $t = 2$ ), they can also be interpreted as uninformed long-term speculators, as in Froot, Scharfstein, and Stein (1992).

<sup>8</sup> Vayanos (2001) studies the strategic trading activity of a risk-averse speculator endowed with a privately observed inventory but no private information about payoffs.

In Kyle (1985) and Pasquariello and Vega (2007), speculators are risk-neutral, hence indifferent to their intermediate wealth and endowment of risky assets. However, we intend to explore the impact of specific endowment shocks on the process of price formation of otherwise identical assets. To that purpose, we further assume that our speculators, albeit risk-neutral, care about the interim as well as the terminal value of their portfolios. Specifically, we assume that each speculator's optimal demands  $x_1(k)$  and  $x_2(k)$  maximize the expected value of the following separable utility function  $U(k)$  of his wealth at  $t = 1$  and 2:

$$U(k) = \gamma W_1(k) + (1 - \gamma)W_2(k), \quad (1)$$

where  $\gamma \in [0,1]$ ,  $W_1(k) = W_0(k) + e_2(k)(p_{1,2} - p_0)$ , and  $W_2(k) = W_0(k) + e_2(k)(v - p_0) + x_1(k)(v - p_{1,1}) + x_2(k)(v - p_{1,2})$ .  $W_1(k)$  is known at the end of the first period, after the MMs set  $p_{1,1}$  and  $p_{1,2}$ , while  $W_2(k)$  is known at the end of the second period, after  $v$  is realized. We interpret the ratio  $\gamma/(1 - \gamma)$  as the speculators' intertemporal marginal rate of substitution between short- and long-term wealth. If  $\gamma = 0$ , each speculator  $k$  reduces to a (long-term) profit-maximizing trader, as in Kyle (1985) and Pasquariello and Vega (2007). If  $\gamma > 0$ , his expected utility at  $t = 1$ , before trading occurs, is given by

$$\begin{aligned} E_1^k[U(k)] &= W_0(k) + \gamma e_2(k)[E_1^k(p_{1,2}) - p_0] \\ &+ (1 - \gamma)\{e_2(k)[E_1^k(v) - p_0] \\ &+ x_1(k)[E_1^k(v) - E_1^k(p_{1,1})] + x_2(k)[E_1^k(v) - E_1^k(p_{1,2})]\}. \end{aligned} \quad (2)$$

At both dates  $t = 1$  and 2 the change in wealth with respect to  $W_0(k)$  depends on two components: the change in value of the existing endowment of asset 2 and the profits from trading in both assets 1 and 2 at  $t = 1$ . However, because the MMs set  $p_{1,1}$  and  $p_{1,2}$  after having observed the corresponding order flow, the value of the net position accumulated at  $t = 1$  is equal to zero in  $W_1(k)$ . This objective function, introduced by Bhattacharya and Nanda (2008) in a single-security framework, can be motivated by wealth constraints, solvency issues, agency and reputation problems, or cash redemptions and injections affecting the interim life of sophisticated market participants such as (open-end) mutual funds. In the context of the U.S. Treasury market, the presence of an active market for security borrowing to deliver against short sales and to avoid settlement fails (the *repo* market) and the possibility of short squeezing when a security is scarce (or *special*) provide additional motivations for why informed traders may care about their interim wealth when trading in the secondary market for government bonds.<sup>11</sup>

### 2.1.2. Equilibrium

Consistently with Kyle (1985), we define a Bayesian Nash equilibrium as a set of  $2(M + 1)$  functions

$x_i(1)(\cdot), \dots, x_i(M)(\cdot)$ , and  $p_{1,i}(\cdot)$  such that the following two conditions hold:

1. *Utility maximization*:  $x_i(k)(\delta_v(k), e_2(k)) = \arg \max E_1^k[U(k)]$ .
2. *Semi-strong market efficiency*:  $p_{1,i} = E(v|\omega_{1,i})$ .<sup>12</sup>

We restrict our attention to linear equilibria. We first conjecture general linear functions for the pricing rule and speculators' demands. We then solve for their parameters satisfying conditions 1 and 2. Finally, we show that these parameters and those functions represent a rational expectations equilibrium. The following proposition accomplishes this task.

**Proposition 1.** *There exists a unique linear equilibrium given by the price functions*

$$p_{1,1} = p_0 + \lambda_1 \omega_{1,1}, \quad (3)$$

$$p_{1,2} = p_0 + \lambda_2 \omega_{1,2}, \quad (4)$$

and by each speculator's demand strategies

$$x_1(k) = \frac{\sigma_z}{\sqrt{M\rho}\sigma_v} \delta_v(k), \quad (5)$$

$$x_2(k) = \frac{\sigma_n}{\sqrt{M\rho}\sigma_v} \delta_v(k) + \frac{1}{2} \frac{\gamma}{1 - \gamma} e_2(k), \quad (6)$$

where  $\sigma_n^2 = \sigma_z^2 + M/4(\gamma/1 - \gamma)^2 \sigma_e^2$ ,  $\lambda_1 = \sqrt{M\rho}\sigma_v/\sigma_z[2 + (M - 1)\rho] > 0$ , and  $\lambda_2 = \sqrt{M\rho}\sigma_v/\sigma_n[2 + (M - 1)\rho] > 0$ .

In equilibrium, each speculator, albeit risk-neutral, exploits his private information cautiously ( $|x_i(k)| < \infty$ ) and in both assets to limit dissipating his informational advantage with his trades. Both optimal trading strategies  $x_i(k)$  depend on his information endowment about the asset payoff ( $\delta_v(k)$ ) and on the corresponding market's depth ( $\lambda_i^{-1}$ ), as in Kyle (1985). Further, as in Pasquariello and Vega (2007), both  $x_1(k)$  (Eq. (5)) and  $x_2(k)$  (Eq. (6)) depend on the number of speculators ( $M$ ) and the correlation among their information endowments ( $\rho$ ). Intuitively, the intensity of competition among speculators affects their ability to maintain the informativeness of the order flow as low as possible. A greater number of speculators trade more aggressively—i.e., their aggregate amount of trading is higher—since (imperfect) competition among them precludes any collusive trading strategy. The heterogeneity of speculators' signals attenuates their trading aggressiveness. When information is less correlated ( $\rho$  closer to zero), each speculator has some monopoly power on his signal, because at least part of it is known exclusively to him. Hence, each speculator trades more cautiously—i.e., his market order is lower—to reveal less of his own information endowment  $\delta_v(k)$ . Thus, either higher  $M$  or  $\rho$  leads to higher equilibrium liquidity in both markets, i.e., lower  $\lambda_1$  and  $\lambda_2$ . This reflects MMs' attempt to be compensated for the losses they anticipate from trading with speculators, as  $\lambda_1$  and  $\lambda_2$  affect their profits from liquidity trading.

<sup>11</sup> For a detailed description of the functioning of the repo market for U.S. Treasury securities see Fleming and Garbade (2007).

<sup>12</sup> Equivalently, competition is assumed to force MMs' expected profits in each market to zero.

### 2.1.3. Testable implications

In the equilibrium of Proposition 1, only the market for asset 2 is affected by the presence of speculators' endowments of that asset and only when their interim wealth ( $W_1(k)$ ) is relevant in their objective function ( $\gamma > 0$ ). When  $\gamma = 0$ , equilibrium speculative trading, liquidity, and prices are the same in both markets:  $x_1(k) = x_2(k) = (\sigma_z / \sqrt{M\rho}\sigma_v)\delta_v(k)$  of Eq. (5) is the optimal informational demand schedule of Kyle (1985) and Pasquariello and Vega (2007),  $\lambda_1 = \lambda_2$ , and  $p_{1,1} = p_{1,2}$ . When  $\gamma > 0$ , such demand schedule remains only in the off-the-run market (asset 1), while the optimal trades in the on-the-run security ( $x_2(k)$ ) also depend on speculators' endowments. This stems from the resolution of a trade-off between short- and long-term wealth: each speculator trades in the on-the-run asset more (or less) than in the off-the-run asset—i.e., more (or less) than he otherwise would if  $\gamma = 0$ —to distort prices in the direction of his endowment shock  $e_2(k)$  and so increase his interim wealth  $W_1(k)$ , regardless of his private signal.<sup>13</sup>

Thus, a portion of each speculator's trade  $x_2(k)$  of Eq. (6) is uninformative about fundamentals ( $v$ ). This in turn implies that the MMs perceive the threat of adverse selection in the market for asset 2 as less serious than in the market for asset 1, so penalize less their counterparts in the former by making it more liquid than the latter:

$$\Delta\lambda = \lambda_1 - \lambda_2 = \frac{\sqrt{M\rho}\sigma_v(\sigma_n - \sigma_z)}{\sigma_z\sigma_n[2 + (M-1)\rho]} > 0, \quad (7)$$

since  $\sigma_z^2 < \sigma_n^2$ . Accordingly, the greater  $\gamma$  and  $\sigma_e$ , the greater is the perceived intensity of uninformative trading in the aggregate order flow for asset 2 (i.e., the greater is  $\sigma_n^2$ ), the less severe is adverse selection for the MMs in that market, thus the greater is the liquidity differential between asset 1 and asset 2. Similarly, greater ex ante uncertainty about both assets' common terminal value  $v$  ( $\sigma_v^2$ ) makes speculators' private information about it more valuable and adverse selection for the MMs in both

markets more severe, yet the less so in the market for asset 2 (where uninformative trading is more intense:  $\sigma_n^2 > \sigma_z^2$ ), thus increasing their liquidity differential. The following corollary summarizes the first set of empirical implications of our model.

**Corollary 1.** *Equilibrium market liquidity in the on-the-run asset is greater than in the off-the-run asset, the more so the greater the relevance of and uncertainty about endowment shocks and the greater the uncertainty about both assets' common fundamentals.*

To gain further insight on the liquidity differential between on-the-run and off-the-run securities, we construct a simple numerical example by setting  $\sigma_v = \sigma_z = \sigma_e = 1$  and  $\gamma = 0.5$ . We then vary the private signal correlation  $\rho$  to study the impact of different degrees of information heterogeneity on the liquidity differential between asset 2 and asset 1 when  $M = 2, 4, 8$ , and 200. We plot the resulting  $\Delta\lambda$  in Fig. 1A. In the presence of numerous speculators (high  $M$ ), the plot for  $\Delta\lambda$  is negatively sloped. Intuitively, more homogeneous private signals (higher  $\rho$ ) attenuate their incentives to behave cautiously when trading. This leads to greater market liquidity in both asset markets, yet the more so in the market for the off-the-run security, where adverse selection is the most severe. Hence, the liquidity differential decreases. However, in the presence of few—thus already less competitive—speculators (low  $M$ ), the plot for  $\Delta\lambda$  is instead positively sloped. Specifically, the equilibrium liquidity differential is lower when those speculators are heterogeneously informed (low  $\rho$ ), since their marginally more cautious use of private information has a smaller impact on their trading activity in the off-the-run market than in the on-the-run market. The following remark formalizes this result.

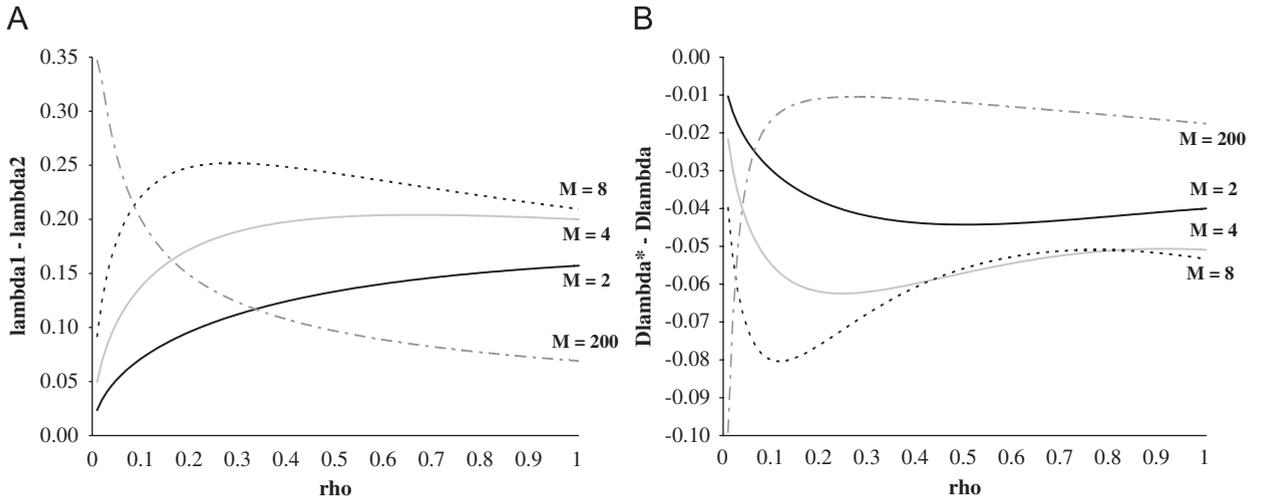
**Remark 1.** In the presence of many (few) speculators, the off/on-the-run liquidity differential is generally increasing (decreasing) in the heterogeneity of their private signals.

## 2.2. Extension: a public signal

The model of Section 2.1 suggests an explanation for the on-the-run liquidity phenomenon that relies on the uncertainty surrounding auction outcomes for just-issued securities. Within this setting, we relate the magnitude of the liquidity differential between cash flow-equivalent assets to the heterogeneity of sophisticated speculators' private signals (and resulting trading activity). To our knowledge, this analysis is novel to the literature. In this section, we investigate the impact of public signals on the on-the-run liquidity phenomenon. Many recent studies investigate the functioning of government bond markets in proximity of the release of macroeconomic news (e.g., Brandt and Kavajecz, 2004; Green, 2004; Pasquariello and Vega, 2007). Yet, the effect of the availability of public information on the relation between on-the-run and off-the-run securities has never been previously explored.

To that purpose, we extend the basic economy by providing each player with an additional, common source of information about the liquidation value of assets 1 and

<sup>13</sup> In equilibrium, these efforts are successful— $\text{cov}(p_{1,2}, e_2(k)) = \frac{1}{2}(\gamma/(1-\gamma))\lambda_2\sigma_e^2 > 0$  and identical to the expected short-term change in the value of each speculator's endowment  $E[e_2(k)(p_{1,2} - p_0)]$ , albeit at the cost of smaller expected long-term profits since  $x_2(k) \neq x_1(k)$ —and may lead to a wedge between off-the-run and on-the-run asset prices  $p_{1,1}$  and  $p_{1,2}$ . This raises the issue of whether the equilibrium of Proposition 1 is consistent with riskless arbitrage exploiting any price difference between two assets with identical liquidation value  $v$ . However, there is no arbitrage opportunity in this model. To show this, we follow Subrahmanyam (1991). First, both  $p_{1,1}$  and  $p_{1,2}$  incorporate all of the individual private signals  $S_v(k)$ ; hence, those prices' unconditional expected difference is zero:  $E(p_{1,1}) = E(p_{1,2})$ . Further, traders in these market neither observe nor can predict with precision the market-clearing prices in assets 1 and 2 when submitting their demands. Thus, profitable arbitrage between those assets is not feasible in equilibrium. Consistently, numerous studies (e.g., Amihud and Mendelson, 1991; Krishnamurthy, 2002; Strebulaev, 2002; Goldreich, Hanke and Nath, 2005) find that on average, after accounting for transaction costs, observed yield differentials between on-the-run and off-the-run securities are too small to represent a viable arbitrage. Along those lines, Eqs. (3) and (4) also imply that  $\text{var}(p_{1,1}) = \text{var}(p_{1,2}) = (M\rho/[2 + (M-1)\rho])\sigma_v^2$ , i.e., speculators' endowment-motivated trading in asset 2 ( $\gamma > 0$ ) does not affect the relative informativeness of that market. Intuitively, more uninformative trading in asset 2 increases informed trading aggressiveness in that market, and hence, does not destabilize its equilibrium price, as in Kyle (1985).



**Fig. 1.** Off/on-the-run liquidity differential without and with a public signal. (A) plots the off/on-the-run liquidity differential  $\Delta\lambda = \lambda_1 - \lambda_2$  is the equilibrium of Proposition 1, in the absence of a public signal of both asset 1 and 2's liquidation value  $v$ , defined in Eq. (7), as a function of  $\rho$ , the degree of correlation of the speculators' private signals, when  $\sigma_v = \sigma_u = \sigma_e = 1$ ,  $\gamma = 0.5$ , and the number of speculators  $M = 2, 4, 8$ , or 200. (B) plots the difference between the off/on-the-run liquidity differential in the presence of a public signal of  $v$  and the differential plotted in (A), i.e.,  $\Delta\lambda^* - \Delta\lambda$  of Eq. (8), again as a function of  $\rho$ , when in addition to the parametrization in (A),  $\sigma_p = 1.5$ .

2 before trading takes place. Specifically, we assume that, sometime between  $t = 0$  and 1, both the speculators and the MMs receive a public and noisy signal  $S_p$  of assets 1 and 2's payoff  $v$ . This signal is normally distributed with mean  $p_0$  and variance  $\sigma_p^2 > \sigma_v^2$ . We further impose that  $cov(S_p, v) = cov[S_p, S_v(k)] = \sigma_v^2$ , so that the parameter  $\sigma_p^2$  controls for the quality of the public signal, and that  $cov[S_p, e_2(k)] = 0$ .

The availability of  $S_p$  affects the level and improves the precision of the information endowments of all market participants prior to trading at time  $t = 1$ , with respect to the economy of Section 2.1. In particular, the MMs' revised beliefs about  $v$  are now given by  $p_0^* = E(v|S_p) = p_0 + (\sigma_v^2/\sigma_p^2)(S_p - p_0)$  and  $\sigma_v^{2*} = var(v|S_p) = \sigma_v^2(1 - \sigma_v^2/\sigma_p^2) < \sigma_v^2$ . The new information endowment of each speculator is  $\delta_v^*(k) = E[v|S_v(k), S_p] - p_0^* = \rho^*[S_v(k) - p_0^*]$ , where  $\rho^* = \rho(\sigma_p^2 - \sigma_v^2)/(\sigma_p^2 - \rho\sigma_v^2) \leq \rho$  is the correlation between any two  $\delta_v^*(k)$  and  $\delta_v^*(j)$ . Hence, we can interpret  $\delta_v^*(k)$  as the truly private (hence less correlated) component of speculator  $k$ 's original private information endowment ( $\delta_v(k)$ ) in the presence of a public signal of  $v$ . The resulting unique linear equilibrium of this amended economy mirrors that of Proposition 1, and is obtained by replacing  $p_0, \sigma_v^2, \rho$ , and  $\delta_v(k)$  with  $p_0^*, \sigma_v^{2*}, \rho^*$ , and  $\delta_v^*(k)$ , respectively, in Eqs. (3) to (6).

2.2.1. Additional testable implications

Pasquariello and Vega (2007) show that, in a Kyle (1985) setting similar to ours, introducing a public signal improves market liquidity. This is the case in our economy as well. Intuitively, the availability of a public signal of  $v$ —by making the speculators' private information less valuable and their trading activity less cautious—reduces the adverse selection risk for the MMs in both the markets

for assets 1 and 2, thus increasing their depth.<sup>14</sup> In this study, we are interested in the impact of the availability of  $S_p$  on the liquidity differentials between on-the-run (2) and off-the-run (1) assets.

To that purpose, we compare the off/on-the-run liquidity differential in the presence of a public signal ( $\Delta\lambda^*$ ) to the one in its absence ( $\Delta\lambda$  in Eq. (7)) as follows:

$$\Delta\lambda^* - \Delta\lambda = \left\{ \frac{(\sigma_p^2 - \sigma_v^2)[2 + (M - 1)\rho]}{\sigma_p \sqrt{\sigma_p^2 - \rho\sigma_v^2}[2 + (M - 1)\rho^*]} - 1 \right\} \Delta\lambda < 0. \tag{8}$$

The availability of a public signal lowers the off/on-the-run liquidity differential since it reduces the perceived adverse selection risk for the MMs in both markets, yet the most in the market for the off-the-run asset (1) where—in absence of endowment-motivated trades—that risk was the greatest in the equilibrium of Proposition 1. This effect is stronger when the available public signal is more precise (lower  $\sigma_p^2$ ), i.e., when the speculators' original private information endowments are less valuable and their trading activity is less cautious.

**Corollary 2.** *The availability of a public signal decreases the off/on-the-run liquidity differential, the more so the lower is that signal's volatility.*

The impact of those endowments' heterogeneity on  $\Delta\lambda^* - \Delta\lambda$  is, however, less obvious, as the following remark illustrates.

<sup>14</sup> It can in fact be shown that  $\lambda_1^* = \sqrt{M\rho^*\sigma_v^2}/\sigma_z[2 + (M - 1)\rho^*] < \lambda_1$  and  $\lambda_2^* = \sqrt{M\rho^*\sigma_v^2}/\sigma_n[2 + (M - 1)\rho^*] < \lambda_2$ , consistent with Pasquariello and Vega (2007).

**Remark 2.** In the presence of many (few) speculators and a public signal, the ensuing decrease in the off/on-the-run liquidity differential is generally increasing (decreasing) in the heterogeneity of their private signals.

In Fig. 1B, we plot  $\Delta\lambda^* - \Delta\lambda$ , the decline in the off/on-the-run liquidity differential due to the availability of a public signal, as a function of  $\rho$ , the correlation of speculators' private signals  $S_p(k)$ , when  $\sigma_p = 1.5$ ,  $\gamma = 0.5$ , and  $M = 2, 4, 8$ , and 200. In the presence of numerous speculators (high  $M$ ), that decline is larger when speculators' private signals are weakly correlated (low  $\rho$ ), since then the impact of  $\rho$  on the aggressiveness of their trading activity is greater, hence, so is the impact of the availability of a public signal on the perceived severity of adverse selection risk in the off-the-run market (asset 1). Fewer speculators (low  $M$ ) already trade more cautiously with their information endowments, and especially so in the off-the-run asset where they suffer no endowment shocks, making that market less liquid (Fig. 1A). Thus, the availability of a public signal reduces the off/on-the-run liquidity differential the most when their incentive to trade cautiously is the lowest (high  $\rho$ ).

### 3. Data description

We test the implications of the model presented in the previous section in a comprehensive sample of U.S. Treasury bond market transaction-level data and U.S. macroeconomic announcements.

#### 3.1. Bond market data

We use intraday, transaction-level data for the most recently issued—*on-the-run*—and the second most recently issued—i.e., just off-the-run—U.S. Treasury securities, consistent with both existing literature (e.g., Krishnamurthy, 2002; Goldreich, Hanke and Nath, 2005) and widespread market practices, between 1992 and 2000. We obtain the data from GovPX, a firm that collects quote and trade information from six of the seven main interdealer brokers (with the notable exception of Cantor Fitzgerald).<sup>15</sup> Fleming (1997) argues that these six brokers account for approximately two-thirds of the voice interdealer-broker market, which in turn translates into approximately 45% of the trading volume in the secondary market for Treasury securities.

In particular, our sample includes every transaction taking place during “regular trading hours,” from 7:30 a.m. to 5:00 p.m. Eastern Time (ET), between January 2, 1992 and December 29, 2000. GovPX stopped recording intraday volume afterward.<sup>16</sup> Strictly speaking, the U.S.

<sup>15</sup> Over our sample period, the major interdealer brokers in the U.S. Treasury market are Cantor Fitzgerald Inc., Garban Ltd., Hilliard Farber & Co. Inc., Liberty Brokerage Inc., RMJ Securities Corp., and Tullet and Tokyo Securities Inc. During that time, Cantor Fitzgerald's share of the interdealer Treasury market—about 30%, according to Goldreich, Hanke and Nath (2005)—came almost exclusively from the “long end” of the Treasury yield curve.

<sup>16</sup> According to Mizraich and Neely (2006), voice-brokered trading volume—the one reported through GovPX—began to decline after 1999

Treasury market is open 24 hours a day; yet, 95% of the trading volume occurs during those hours (e.g., Fleming, 1997). Thus, to remove fluctuations in bond prices due to illiquidity, we ignore trades outside that narrower interval. We analyze three-month, six-month, and one-year Treasury bills and two-year, five-year, and 10-year Treasury notes. Not included in the analysis are three-year notes (because the U.S. Treasury suspended their issuance in 1998), 30-year bonds (because of limited sample coverage by GovPX, see Footnote 15), and Treasury inflation-indexed securities (because of their limited trading activity over the sample period), as in Fleming (2003). Finally, the data contain some interdealer brokers' posting errors not previously filtered out by GovPX. We eliminate these errors following the procedure described in Fleming (2003).

We complement the GovPX data with information on those bills' and notes' fundamental characteristics (daily modified duration and convexity) from Morgan Markets, and with official data on the history of those bonds' routinely scheduled Treasury auctions: the date of the auction, the amount of competitive, non-competitive, and System Open Market Account (SOMA) tenders ( $Ten_t$ , a measure of government debt demand), the amount of tenders accepted by the U.S. Treasury ( $Acc_t$ , a measure of government debt supply), and high, low, and average accepted competitive yield bids. This information is publicly available on the U.S. Treasury Web site.<sup>17</sup>

We report summary statistics for the following variables in Table 1A (on-the-run and off-the-run Treasury bills) and Table 1B (on-the-run and off-the-run Treasury notes): average daily quoted percentage bid–ask spreads ( $S_t^{on}$  and  $S_t^{off}$ ), modified duration ( $D_t^{on}$  and  $D_t^{off}$ ), modified convexity ( $C_t^{on}$  and  $C_t^{off}$ ), total amount tendered at the auction, total amount accepted at the auction, and range of competitive yield bids at the auction (highest bid minus lowest bid divided by average accepted competitive bid,  $HL_t$ ). Consistent with market conventions (e.g., see Fleming, 2003), Treasury bills in our sample are quoted in terms of a discount rate, while Treasury notes are quoted in points, i.e., as a percentage of par. In both instances, percentage bid–ask spreads are computed as a fraction of the (discount or price) midquote multiplied by 100; total amounts tendered and accepted are in billions of U.S. dollars; modified durations are in fractions of 365 days.

#### 3.2. Macroeconomic data

The model of Section 2 relates the off/on-the-run liquidity differential to the heterogeneity of private information about fundamentals among sophisticated market participants, as well as to the release of public information about those fundamentals. In this paper, we

(footnote continued)

as electronic trading platforms (e.g., eSpeed, BrokerTec) became available. Those platforms now account for most Treasury market trading activity. Our evidence is nevertheless robust to removing all transactions occurring in 2000 from our sample.

<sup>17</sup> <http://www.publicdebt.treas.gov>.

**Table 1A**

U.S. Treasury bills: summary statistics.

This table presents the mean and standard deviation for several variables used in our empirical tests for three-month, six-month, and one-year on-the-run and just off-the-run Treasury bills between January 2, 1992 and December 29, 2000.  $S_t$  is the average daily bid–ask percentage discount spread, computed as the daily average of 100 times the difference between the intraday offer and bid discount rates divided by the corresponding discount midquote (from GovPX, a firm that collects quote and trade information from six of the seven main Treasury interdealer brokers).  $C_t$  and  $D_t$  are the daily convexity and modified duration (from Morgan Markets).  $Ten_t$  and  $Acc_t$  are the total amounts tendered and accepted at the corresponding U.S. Treasury auctions, in billions of U.S. dollars.  $HL_t$  is the highest competitive bid minus the lowest competitive bid divided by the average competitive bid at the corresponding U.S. Treasury auctions. Auction data is from the U.S. Treasury. The table also reports the difference between the means of  $S_t$ ,  $D_t$ , and  $C_t$  for the corresponding just off-the-run and on-the-run Treasury bills. A \*, \*\*, or \*\*\* indicates statistical significance at the 10%, 5%, or 1% levels, respectively.

	Off-the-run		On-the-run		Difference in Mean
	Mean	Stdev.	Mean	Stdev.	
Three-month					
Bid–ask discount spread: $S_t$	0.291	0.240	0.120	0.067	0.171***
Convexity: $C_t$	0.0010	0.0001	0.0012	0.0001	–0.0002***
Modified duration: $D_t$	0.221	0.007	0.240	0.009	–0.019***
Total amount tendered: $Ten_t$			40.850	10.867	
Total amount accepted: $Acc_t$			12.308	1.689	
Range of competitive bids: $HL_t$			0.005	0.003	
Six-month					
Bid–ask discount spread: $S_t$	0.260	0.172	0.130	0.057	0.130***
Convexity: $C_t$	0.0041	0.0005	0.0045	0.0004	–0.0004***
Modified duration: $D_t$	0.452	0.030	0.473	0.027	–0.021***
Total amount tendered: $Ten_t$			38.924	10.185	
Total amount accepted: $Acc_t$			12.332	1.277	
Range of competitive bids: $HL_t$			0.004	0.002	
One-year					
Bid–ask discount spread: $S_t$	0.275	0.168	0.110	0.047	0.165***
Convexity: $C_t$	0.010	0.003	0.012	0.003	–0.002***
Modified duration: $D_t$	0.789	0.164	0.892	0.147	–0.102***
Total amount tendered: $Ten_t$			47.086	11.855	
Total amount accepted: $Acc_t$			17.266	2.056	
Range of competitive bids: $HL_t$			0.004	0.002	

**Table 1B**

U.S. Treasury notes: summary statistics.

This table presents the mean and standard deviation of several variables used in our empirical tests for two-year, five-year, and 10-year on-the-run and just off-the-run Treasury notes between January 2, 1992 and December 29, 2000.  $S_t$  is the average daily bid–ask percentage price spread, computed as the daily average of 100 times the difference between the intraday offer and bid prices divided by the corresponding price midquote (from GovPX, a firm that collects quote and trade information from six of the seven main Treasury interdealer brokers).  $C_t$  and  $D_t$  are the daily convexity and modified duration (from Morgan Markets).  $Ten_t$  and  $Acc_t$  are the total amounts tendered and accepted at the corresponding U.S. Treasury auctions, in billions of U.S. dollars.  $HL_t$  is the highest competitive bid minus the lowest competitive bid divided by the average competitive bid at the corresponding U.S. Treasury auctions. Auction data is from the U.S. Treasury. The table also reports the difference between the means of  $S_t$ ,  $D_t$ , and  $C_t$  for the corresponding just off-the-run and on-the-run Treasury notes. A \*, \*\*, or \*\*\* indicates statistical significance at the 10%, 5%, or 1% levels, respectively.

	Off-the-run		On-the-run		Difference in Mean
	Mean	Stdev.	Mean	Stdev.	
Two-year					
Bid–ask price spread: $S_t$	0.017	0.005	0.008	0.002	0.009***
Convexity: $C_t$	0.040	0.001	0.044	0.001	–0.003***
Modified duration: $D_t$	1.763	0.035	1.842	0.034	–0.080***
Total amount tendered: $Ten_t$			41.802	6.341	
Total amount accepted: $Acc_t$			18.300	1.844	
Range of competitive bids: $HL_t$			0.013	0.006	
Five-year					
Bid–ask price spread: $S_t$	0.030	0.009	0.014	0.004	0.016***
Convexity: $C_t$	0.204	0.010	0.214	0.007	–0.010***

Table 1B (continued)

	Off-the-run		On-the-run		Difference in Mean
	Mean	Stdev.	Mean	Stdev.	
Modified duration: $D_t$	4.110	0.113	4.223	0.087	-0.113***
Total amount tendered: $Ten_t$			30.679	3.736	
Total amount accepted: $Acc_t$			12.914	1.830	
Range of competitive bids: $HL_t$			0.015	0.018	
10-year					
Bid-ask price spread: $S_t$	0.054	0.014	0.024	0.005	0.030***
Convexity: $C_t$	0.596	0.027	0.641	0.033	-0.045***
Modified duration: $D_t$	6.824	0.203	7.106	0.244	-0.283***
Total amount tendered: $Ten_t$			30.676	4.244	
Total amount accepted: $Acc_t$			13.385	2.093	
Range of competitive bids: $HL_t$			0.008	0.018	

Table 2

Dispersion of beliefs: summary statistics.

This table presents summary statistics for the standard deviation across professional forecasts for 18 macroeconomic news announcements (from Money Market Services (MMS)) over the full sample period between January 1992 and December 2000. We report the number of observations, mean, standard deviation, maximum, minimum, Spearman rank correlation with the Nonfarm Payroll standard deviation,  $\rho(\text{Payroll})$ , and first-order autocorrelation coefficient,  $\rho(1)$ . A \*, \*\*, or \*\*\* indicates statistical significance at 10%, 5%, or 1% levels, respectively.

	Obs.	Mean	Stdev.	Max	Min	$\rho(\text{Payroll})$	$\rho(1)$
Quarterly announcements							
1. GDP advance	36	0.480	0.170	1.100	0.320	0.162*	-0.181
2. GDP preliminary	34	0.313	0.178	1.290	0.120	0.014	0.192
3. GDP final	35	0.128	0.051	0.240	0.040	0.083	0.250
Monthly announcements							
<i>Real activity</i>							
4. Nonfarm payroll	108	41.814	14.212	103.190	17.496	1.000	0.424***
5. Retail sales	107	0.302	0.158	1.390	0.106	0.109	0.047
6. Industrial production	107	0.183	0.135	1.700	0.087	0.236**	0.358***
<i>Consumption</i>							
7. New home sales	106	19.270	10.235	96.225	7.840	0.151	0.099
<i>Investment</i>							
8. Durable goods orders	106	1.034	0.333	2.583	0.450	0.077	0.412***
9. Factory orders	105	0.587	0.577	7.249	0.239	0.219**	0.015
10. Construction spending	105	0.499	0.253	1.270	0.158	0.176*	0.192***
<i>Net exports</i>							
11. Trade balance	107	0.790	0.851	11.480	0.400	0.122	0.018
<i>Prices</i>							
12. Producer price index	108	0.130	0.049	0.380	0.060	0.186*	0.287***
13. Consumer price index	107	0.086	0.051	0.580	0.040	0.146	0.221**
<i>Forward-looking</i>							
14. Consumer confidence index	106	1.646	0.609	4.026	0.663	0.079	0.230**
15. NAPM index	107	0.961	0.303	2.680	0.441	0.242**	0.382***
16. Housing starts	106	0.045	0.038	0.430	0.016	0.160	0.246***
17. Index of leading indicators	108	0.202	0.137	0.920	0.044	0.134	0.480***
Weekly announcements							
18. Initial unemployment claims	459	7.973	5.440	53.400	2.100	0.069	0.578***

use the International Money Market Services (MMS) Inc. real-time data on the release dates and professional forecasts of 25 of the most relevant U.S. macroeconomic announcements.<sup>18</sup> We use the MMS standard deviation across those forecasts as a measure of the dispersion of beliefs across speculators. This measure of information heterogeneity is widely adopted in the literature on investors' reaction to information releases in the stock market (e.g., Diether, Malloy, and Scherbina, 2002; Kallberg and Pasquariello, 2008); Green (2004) and Pasquariello and Vega (2007) recently use it in the bond market. The 18 macroeconomic announcements for which this variable is available in our sample, the corresponding number of observations, and the reporting agency are listed in Table 2.

The dispersion of beliefs is positively correlated across the macroeconomic announcements in our sample, yet not strongly so. For instance, Pasquariello and Vega (2007) report that the pairwise correlation between each announcement and arguably the most important of them, the Nonfarm Payroll report (e.g., Andersen and Bollerslev, 1998; Andersen, Bollerslev, Diebold, and Vega, 2007; Brenner, Pasquariello, Subrahmanyam, 2009), is positive, albeit not statistically significant for most of the announcements in the sample ( $\rho(\text{Payroll})$  in Table 2). Thus, we follow Pasquariello and Vega (2007) and construct three alternative measures of dispersion of beliefs during announcement and non-announcement days: one based exclusively on the Payroll announcement, another based on seven "influential" announcements (Nonfarm Payroll Employment, Retail Sales, New Home Sales, Consumer Confidence Index, NAPM Index, Index of Leading Indicators, and Initial Unemployment Claims), and the last one based on the 18 announcements in Table 2. Pasquariello and Vega (2007) label the seven macroeconomic announcements listed above "influential" for they are the only ones having a statistically significant impact on day-to-day bond yield changes over our sample period.

We then define a monthly proxy for the aggregate degree of information heterogeneity about macroeconomic fundamentals as a weighted sum of monthly dispersions across announcements,

$$SSD_{Pt} = \sum_{j=1}^P \frac{SD_{jt} - \hat{\mu}(SD_{jt})}{\hat{\sigma}(SD_{jt})}, \quad (9)$$

where  $SD_{jt}$  is the standard deviation of announcement  $j$  across professional forecasts,  $\hat{\mu}(SD_{jt})$  and  $\hat{\sigma}(SD_{jt})$  are its sample mean and standard deviation, respectively, and  $P$  is equal to either 1 (Nonfarm Payroll Employment), 7 (the "influential" announcements listed above), or 18 (i.e., those in Table 2). The standardization in Eq. (9) is necessary because units of measurement differ across announcements. We use the monthly dispersion estimates from these three methodologies to classify days in

which the corresponding monthly variable  $SSD_{Pt}$  is above (below) the top (bottom) 70th (30th) percentile of its empirical distribution as days with high (low) information heterogeneity. The resulting time series of high (+1) and low (−1) dispersion days are positively correlated: their correlations range from 0.37 (between the Payroll-based series,  $P = 1$ , and the series constructed with the influential announcements,  $P = 7$ ) to 0.70 (between the series using all announcements,  $P = 18$ , and the one based only on the influential news releases,  $P = 7$ ).

#### 4. Empirical analysis

The model of Section 2 generates several implications for the on-the-run liquidity phenomenon in bond markets that we now test in this section. To that purpose, we need to compute off/on-the-run liquidity differentials that are compatible with those in the model for each of the bills and notes in our sample. This is a challenging task. In the context of our model, and consistent with Kyle (1985), market liquidity for a traded asset  $i$  is defined as the marginal impact of an unexpected trade on the equilibrium price of that asset,  $\lambda_i$ . This measure of liquidity is typically estimated as the slope  $\lambda_t$  of the regression of yield or price changes on the observed aggregate order flow (net volume) over either intraday or daily time intervals. Hence, whenever transaction data are available, this procedure allows for a direct assessment of our model's implications for off/on-the-run liquidity differentials  $\Delta\lambda$ .

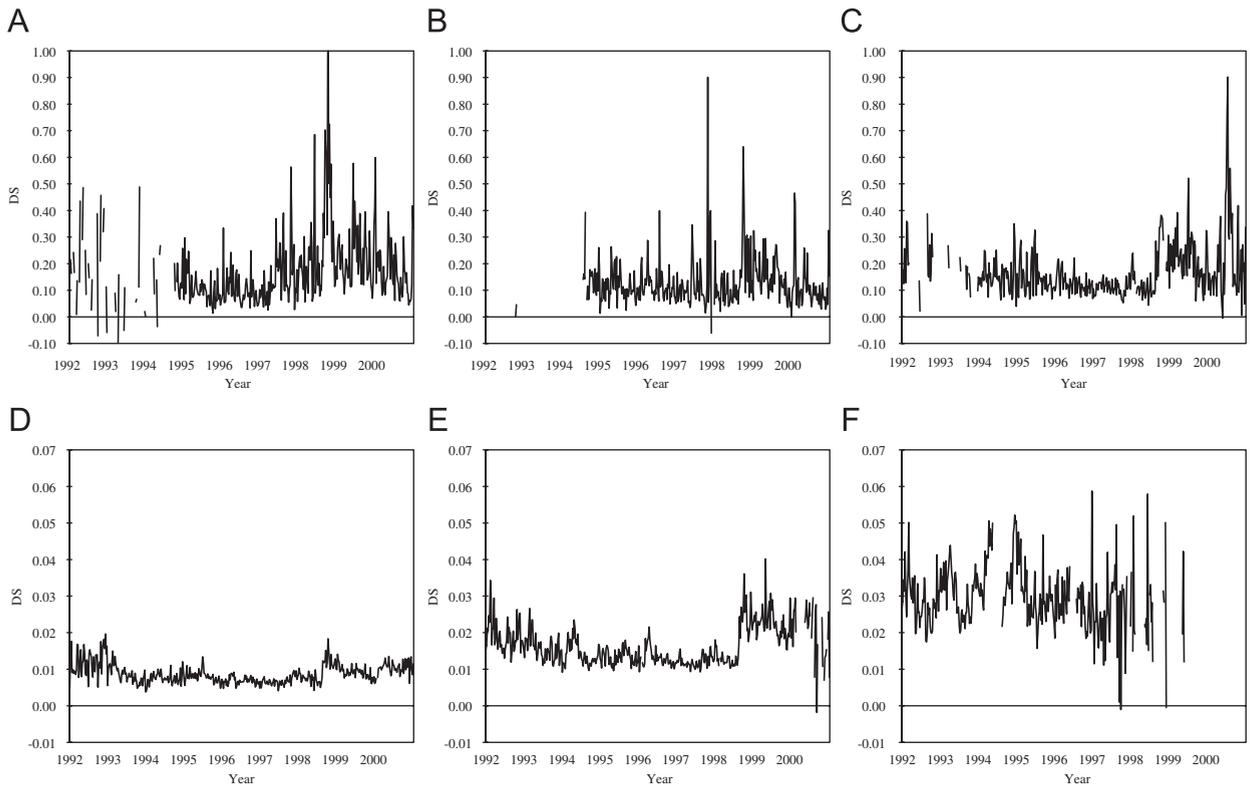
The GovPX database contains such data, i.e., in theory, it allows for the direct estimation of  $\lambda_t^{off}$  and  $\lambda_t^{on}$ . Unfortunately, the relative scarcity of trades (but not of posted bid and ask quotes) in off-the-run bonds often makes the estimation of  $\lambda_t^{off}$  at the daily frequency problematic.<sup>19</sup> In addition, even when possible, the direct estimation of  $\lambda_t$  also suffers from several shortcomings. In particular, it requires the econometrician (i) to specify a model for the prior estimation of the unobserved portion of the aggregate order flow, as well as (ii) to control for many additional microstructure imperfections that, together with informed and liquidity trading, may affect its dynamics (e.g., Hasbrouck, 2007).<sup>20</sup> Thus, any inference from such an effort is subject to potential misspecification, as well as to the potential biases stemming from measurement errors in the dependent variable. The latter are likely to be severe if any independent variable explaining  $\lambda_t$  is also not measured properly (e.g., see the discussion in Greene, 1997, p. 436).

In light of these considerations, in this paper we measure each market's liquidity using its daily average quoted percentage bid–ask discount (for bills) or price (for notes) spread,  $S_t^{on}$  and  $S_t^{off}$  of Section 3.1, for several reasons. First, off-the-run and on-the-run spreads are

<sup>18</sup> Detailed discussions of the properties of this database can be found in Fleming and Remolona (1997), Andersen, Bollerslev, Diebold, and Vega (2003), and Pasquariello and Vega (2007). MMS was acquired by Informa in 2003 and no longer exists; Action Economics is now providing similar survey services.

<sup>19</sup> For instance, Fabozzi and Fleming (2004) report that off-the-run securities account for only 24% of interdealer trading volume in the U.S. Treasury market in 1998.

<sup>20</sup> Green (2004), Brandt and Kavajecz (2004), and Pasquariello and Vega (2007) are recent examples of such efforts in the U.S. Treasury market.



**Fig. 2.** Historical off/on-the-run bid-ask spread differentials. In this figure, we plot weekly averages of daily average off/on-the-run percentage bid-ask discount (for bills) or price (for notes) spread differentials multiplied by 100 ( $\Delta S_t = S_t^{\text{off}} - S_t^{\text{on}}$ , solid line, left axis) over the sample period 1/1992–12/2000 for three-month (A), six-month (B), and one-year (C) U.S. Treasury bills, and two-year (D), five-year (E), and 10-year (F) U.S. Treasury notes.

virtually without measurement error. Further, there is an extensive literature relating their magnitude and dynamics to the informational role of trading (see O'Hara, 1995, for a review). Lastly, when comparing several alternative measures of liquidity in the U.S. Treasury market, Fleming (2003) finds that the quoted bid-ask spread is the most highly correlated with both direct estimates of price impact and well-known episodes of poor liquidity in those markets.<sup>21</sup> The inference that follows is nonetheless robust to replacing  $S_t^{\text{on}}$  and  $S_t^{\text{off}}$  with  $\lambda_t^{\text{on}}$  and  $\lambda_t^{\text{off}}$ , respectively, whenever direct estimation of  $\lambda_t^{\text{off}}$  is feasible, as well as with average daily effective

percentage (discount or price) bid-ask spreads and percentage bid-ask yield spreads.

#### 4.1. The benchmark on-the-run liquidity phenomenon

The main objective of this paper is to study the informational role of bond trading in the presence of auction-driven endowment shocks in explaining off/on-the-run liquidity differentials in fixed income markets. To that purpose, we start by computing average daily off/on-the-run bid-ask spread differentials as  $\Delta S_t = S_t^{\text{off}} - S_t^{\text{on}}$  for three-month, six-month, one-year, two-year, five-year, and 10-year Treasury bills and notes between 1992 and 2000. We then plot the resulting time series of  $\Delta S_t$  in Fig. 2 by week to smooth daily variability, as in Fleming (2003), and report their sample averages in Table 1. There is clear, economically significant evidence of the on-the-run liquidity phenomenon in the U.S. Treasury bond market between 1992 and 2000: off/on-the-run spread differentials  $\Delta S_t$  are large (e.g., on average, never less than 50% of the corresponding off-the-run spreads), always positive (solid line in Fig. 2A–F), and statistically significant across all maturities (at the 1% level, in Tables 1A and 1B). Mean daily spread differentials range from less than one basis point (two-year, Fig. 2D) to more than 17

<sup>21</sup> See also Chordia, Sarkar, and Subrahmanyam (2005) and Goldreich, Hanke and Nath (2005). The relative scarcity of quotes and trades in the off-the-run market mentioned above precludes us from pursuing any of the techniques available in the literature to separate the portion of the bid-ask spread due to adverse selection from those due to order processing costs or inventory control (e.g., Stoll, 1989; George, Kaul, and Nimalendran, 1991). In any case, execution costs are likely to be similar across Treasury bonds, hence to cancel out when computing average daily off/on-the-run spread differentials  $\Delta S_t$ . Furthermore, we find (and discuss in Section 4.1) that those differentials are insensitive to the corresponding repo rate differentials, which may proxy for the relative cost of unwinding undesired inventory positions in the off-the-run and on-the-run markets.

**Table 3**

Benchmark on-the-run liquidity phenomenon.

Panel A reports ordinary least squares (OLS) estimates of the following regression model (Eq. (10)) over the full sample (1/1992–12/2000):

$$\Delta S_t = \beta_0 + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \varepsilon_t,$$

where  $\Delta S_t = S_t^{off} - S_t^{on}$  is the daily average off/on-the-run percentage bid–ask discount (for U.S. Treasury bills) or price (for U.S. Treasury notes) spread differential multiplied by 100,  $\Delta D_t = D_t^{off} - D_t^{on}$  is the off/on-the-run modified duration differential, and  $\Delta C_t = C_t^{off} - C_t^{on}$  is the off/on-the-run convexity differential, as well as the sample mean of  $\Delta S_t$  from Table 1.  $R_a^2$  is the adjusted  $R^2$  from the estimation of the fully specified regression above;  $n$  is the number of observations;  $n_p$  is the number of security pairs;  $\bar{\beta}_0$  is the average security-pair fixed effect from the estimation of Eq. (10) when allowing for both security-pair fixed effects and year dummies. Panel B reports OLS estimates on an amended specification of Eq. (10) including an additional explanatory variable, off/on-the-run repo rate differentials ( $\Delta repo_t$ ), over the subsample for which that variable is available (7/1997–12/2000). A \*, \*\*, or \*\*\* indicates statistical significance at the 10%, 5%, or 1% levels, respectively, using Newey–West standard errors for  $\beta_0$ ,  $\beta_1$ ,  $\beta_2$ , and  $\beta_{\Delta repo}$ .

	Mean $\Delta S_t$	$\beta_0$	$\beta_1$	$\beta_2$	$\beta_{\Delta repo}$	$R_a^2$	$n$	$n_p$	$\bar{\beta}_0$
Panel A: Full sample (1/1992–12/2000)									
Three-month	0.1708***	0.2034***	0.0220	−0.5002		0.12%	1,578	316	0.1240***
Six-month	0.1297***	0.1375***	−0.0287	1.7725		−0.07%	1,350	298	0.1786***
One-year	0.1646***	0.1794***	−0.0113***	0.5470***		2.81%	1,771	94	0.1655***
Two-year	0.0087***	0.0040	0.0005	−0.0267		1.13%	2,197	107	0.0055***
Five-year	0.0161***	0.0126***	0.0005**	−0.0095***		14.80%	2,023	86	0.0139***
10-year	0.0298***	0.0362***	0.0004***	−0.0014*		6.18%	1,266	21	0.0513***
Panel B: Repo sample (7/1997–12/2000)									
Three-month	0.2206***	0.2867***	0.0864	−5.4889	0.0207	−0.11%	834	177	0.2167***
Six-month	0.1412***	0.1725***	−0.1810*	10.6462*	0.0286	0.66%	770	169	0.2104***
One-year	0.1900***	0.1632***	−0.0194***	0.7369**	0.0557	6.20%	811	38	0.1220***
Two-year	0.0093***	−0.0001	−0.0006	−0.0136	0.0010***	7.63%	827	41	0.0051***
Five-year	0.0187***	0.0130***	0.0003	−0.0063**	0.0013***	17.35%	662	20	0.0213***
10-year	0.0251***	0.0320***	−0.0002	0.0032	−0.0006	0.06%	98	5	−0.0430***

basis points (three-month, Fig. 2A). Notably, these graphs reveal occasional gaps in GovPX market coverage, especially among six-month bills and 10-year notes in the earlier and latter parts of the sample period, respectively.<sup>22</sup>

Next, we establish the robustness of the on-the-run liquidity phenomenon in the U.S. Treasury market as reported in Table 1. This is a necessary step in our analysis, for recent studies (e.g., Krishnamurthy, 2002; Strebulaev, 2002; Goldreich, Hanke and Nath, 2005) argue that off/on-the-run yield differentials may either disappear or considerably diminish once controlling for these bonds' fundamental characteristics.<sup>23</sup> Some of those fundamental characteristics are in fact likely to differ for on-the-run bonds and their closest off-the-run securities, although these securities' liquidation values are assumed to be identical in our model. In particular, Table 1 suggests that duration and convexity differentials between them may be large. For instance, both off/on-the-run modified duration and convexity differentials ( $\Delta D_t = D_t^{off} - D_t^{on}$

and  $\Delta C_t = C_t^{off} - C_t^{on}$ , respectively) are always negative and significant at the 1% level. Hence, on-the-run bonds are on average less sensitive to parallel shifts of the yield curve and to large, sudden yield jumps than corresponding off-the-run securities at each maturity. Investors' expectations and risk aversion may then affect their relative preferences toward these assets, i.e., may ultimately affect these assets' relative liquidity in a systematic fashion.

To assess the empirical relevance of these considerations, we specify the following benchmark model of off/on-the-run bid–ask spread differentials:

$$\Delta S_t = \beta_0 + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \varepsilon_t, \tag{10}$$

for each of the bills and notes in our database. We estimate these regressions for each maturity separately by ordinary least squares (OLS) and evaluate the statistical significance of the coefficients' estimates, reported in Panel A of Table 3, with Newey–West standard errors to correct for heteroskedasticity and serial correlation.<sup>24</sup> The time series  $\Delta S_t$ ,  $\Delta D_t$ , and  $\Delta C_t$  are effectively made of several different pairs of first off-the-run and on-the-run securities stacked on each other over the sample period

<sup>22</sup> For a discussion of the incompleteness of GovPX coverage of the U.S. Treasury market, see Boni and Leach (2002) and Fleming (2003).

<sup>23</sup> For example, Goldreich, Hanke and Nath (2005) find that, after adjusting for coupon and maturity differentials with prices of hypothetical Treasury notes, the resulting average daily two-year off/on-the-run yield differential between 1994 and 2000 is small (i.e., never larger than 1.5 basis points at its peak) and rapidly declining to zero during the monthly auction cycle until a newer note is issued (Fig. 2, p. 13).

<sup>24</sup> Our dependent variable  $\Delta S_t$  is serially correlated, although mildly so. Unreported analysis indicates that the serial correlation in off/on-the-run bid–ask spread differentials  $\Delta S_t$  is far from a unit root: the first-order autocorrelation ranges from 0.257 to 0.600 and dies off quickly (e.g., the second-order autocorrelation ranges from 0.010 to 0.300).

(as in Brandt and Kavajecz, 2004; Green, 2004; Pasquariello and Vega, 2007 when studying yield dynamics).<sup>25</sup> Thus, as a robustness check, we also estimate Eq. (10) allowing for security-pair fixed effects and year dummies. The results in Table 3 provide further, strong evidence of the on-the-run liquidity phenomenon in the U.S. Treasury market. For all maturities, both the magnitude and significance of estimates for the average off/on-the-run liquidity differentials ( $\beta_0$  in Eq. (10) or the average security-pair fixed effect  $\bar{\beta}_0$ ) are virtually unaffected—or even amplified—by the inclusion of duration and convexity differentials. We obtain similar results (not reported here) when replacing  $\Delta D_t$  and  $\Delta C_t$  in Eq. (10) with  $D_t^{off}$ ,  $D_t^{on}$ ,  $C_t^{off}$ , and  $C_t^{on}$ , or with coupon and time-to-maturity differentials, as well as controlling for lagged values of  $\Delta S_t$ .

Existing research also suggests that government bond yields and off/on-the-run yield differentials may be related to the relative degree of specialness of these securities as a repo collateral (e.g., Duffie, 1996; Krishnamurthy, 2002; Sundaresan and Wang, 2009; Fleming and Garbade, 2007) or in response to (actual or expected) short squeezes in the post-auction on-the-run market stemming from pre-auction trading and auction allocations (e.g., Nyborg and Sundaresan, 1996; Bindseil, Nyborg, and Strebulaev, 2002; Nyborg and Strebulaev, 2004). Such specialness may in turn affect investors' preferences toward on-the-run and off-the-run securities, thus ultimately their relative liquidity in the secondary market.<sup>26</sup> Alternatively, Amihud and Mendelson (1991) observe that off-the-run securities are more likely to be “locked away” in investors' portfolios, either precluding or obstructing dealers' efforts to supply immediacy—in the sense of Grossman and Miller (1988)—in that market (see also Garbade and Silber, 1976). The ensuing opportunity and inventory-management costs are then likely to be reflected in the repo market (e.g., Krishnamurthy, 2002). We account for the role of the repo market, repo specialness, and short squeezes for the on-the-run liquidity phenomenon by using the (limited) information on repo rates provided by Morgan Markets only from July 1997 onward. Comprehensive data on repo rates for all the securities in our sample and over its entire length is unavailable to us. We amend Eq. (10) to include off/on-the-run repo rate differentials ( $\Delta repo_t$ ), when available, as proxies for the perceived relative scarcity of either security in the secondary market.<sup>27</sup> Despite the more limited sample coverage, the estimation of these amended regressions

<sup>25</sup> Unreported analysis shows that higher-order autocorrelation in  $\Delta S_t$  does not increase around the time when security pairs are replaced (approximately every five business days for bills and 20 business days for bonds). Therefore, the stacked time series  $\Delta S_t$  appear to behave like *continuous* time series, rather than like stacked time series of identical security pairs.

<sup>26</sup> For instance, Graveline and McBrady (2006) provide evidence that repo differentials in the U.S. Treasury market may be more closely related to intermediaries' demand for on-the-run securities to hedge interest rate risk in their inventories.

<sup>27</sup> We thank Arvind Krishnamurthy for recommending this line of action to us. Qualitatively similar inference can be drawn from replacing  $\Delta repo_t$  with the differential between *general collateral* rates and the corresponding on-the-run repo rates.

(reported in Panel B of Table 3) suggests that the impact of those repo differentials on  $\Delta S_t$  (i.e., the coefficient  $\beta_{\Delta repo}$ ) is in most cases statistically insignificant and that our analysis is robust to their inclusion.<sup>28</sup>

Lastly, we consider the possibility that the average off/on-the-run liquidity differentials reported in Table 3 ( $\beta_0$ ) may be biased by term premiums, i.e., vintage effects on the yield curve not captured by differentials in modified duration, convexity, and repo rates. We investigate the importance of this concern by augmenting Eq. (10) with such customary proxies for the slope of the U.S. Treasury yield curve as the end-of-day yield differential between on-the-run (or off-the-run) 10-year notes and three-month bills or the second principal component of the correlation matrix of the end-of-day yields of all the on-the-run securities in the sample, as well as with interaction terms of either proxy and both duration and convexity differentials between off-the-run and off-the-run securities. The latter allow for both the impact of  $\Delta D_t$  and  $\Delta C_t$  on  $\Delta S_t$  to change in response to fluctuations in the slope of the yield curve over time and the impact of term premiums on  $\Delta S_t$  to change in response to fluctuations in  $\Delta D_t$  and  $\Delta C_t$  over time. Details of these additional estimations are available on request. We find the evidence reported in this paper to be virtually unaffected in either specification.<sup>29</sup>

Overall, the results in Tables 1 and 3 and Fig. 2 indicate that off/on-the-run bid-ask spread differentials are positive, economically significant—averaging more than half of the corresponding mean off-the-run spread—and cannot be explained by differences in the fundamental characteristics of the underlying securities.

#### 4.2. Endowment shocks

The analysis so far reveals that (i) off/on-the-run liquidity differentials in the U.S. Treasury market are both economically and statistically significant, and (ii) this phenomenon is not explained away by differences in the underlying securities' fundamentals. We are now ready to test directly the model's main implication for these results, namely that off/on-the-run liquidity differentials

<sup>28</sup> As an anonymous referee points out, if short squeezing played an important role for the on-the-run liquidity phenomenon the distribution of off/on-the-run liquidity differentials in the post-auction market would be bimodal, since an on-the-run security has either great scarcity value or it has no scarcity value. Yet, in unreported analysis, the *dip* test of Hartigan and Hartigan (1985) fails to reject the null hypothesis of unimodality for each of the off/on-the-run bid-ask spread differentials  $\Delta S_t$  in our sample. As a further robustness test, we also consider the impact of changing auction rules over our sample period (see Section 2.1) on our inference, since Nyborg and Strebulaev (2004) show that discriminatory auctions lead to more short squeezing than uniform auctions. Until the early 1990s, the U.S. Treasury employed exclusively discriminatory auctions to sell its securities to the public. However, it began auctioning two-year and five-year notes with the uniform-price format in September 1992; this format was then extended to all other Treasury securities only in October 1998. Nevertheless, in unreported analysis we find no systematic differences in the economic and statistical significance of off/on-the-run bid-ask spread differentials in the pre- and post-uniform auction periods.

<sup>29</sup> Estimating the latent slope factor by spanning the space of maturities in our sample with both on-the-run and off-the-run securities leads to the same conclusion.

**Table 4**

Auction effect: off/on-the-run bid–ask spread differential.

This table reports ordinary least squares (OLS) estimates of the following regression model (Eq. (11)) over the full sample (1/1992–12/2000):

$$\Delta S_t = \beta_0 + \sum_{i=1}^N \beta_{Ai} \text{Auction}_{t-i} + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \varepsilon_t,$$

where  $\Delta S_t = S_t^{\text{off}} - S_t^{\text{on}}$  is the daily average off/on-the-run percentage bid–ask discount (for U.S. Treasury bills) or price (for U.S. Treasury notes) spread differential multiplied by 100,  $\Delta D_t = D_t^{\text{off}} - D_t^{\text{on}}$  is the off/on-the-run modified duration differential, and  $\Delta C_t = C_t^{\text{off}} - C_t^{\text{on}}$  is the off/on-the-run convexity differential,  $\text{Auction}_{t-i}$  is a dummy variable equal to one on day  $t$  if day  $t - i$  is the most recent auction date for the corresponding bond and equal to zero otherwise. We assume  $N = 4$  for three-month and six-month bills and  $N = 10$  for all other bonds.  $R_a^2$  is the adjusted  $R^2$  from the estimation of the fully specified regression above. A \*, \*\*, or \*\*\* indicates statistical significance at the 10%, 5%, or 1 levels, respectively, using Newey–West standard errors.

	Three-month	Six-month	One-year	Two-year	Five-year	10-year
$\beta_0$	0.2521***	0.1499***	0.1970***	0.0064**	0.0145***	0.0365***
$\beta_{A1}$	-0.0894***	-0.057***	-0.0750***	-0.004***	-0.0051***	-0.0105***
$\beta_{A2}$	-0.0514***	-0.0264***	-0.0714***	-0.0024***	-0.0035***	-0.0048***
$\beta_{A3}$	0.0033	-0.0073	-0.053***	-0.0020***	-0.0046***	-0.0085***
$\beta_{A4}$	0.045***	0.0312***	-0.0595***	-0.0028***	-0.0039***	-0.0017
$\beta_{A5}$			-0.0472***	-0.0023***	-0.0033***	-0.0005
$\beta_{A6}$			-0.0376***	-0.0021***	-0.0024***	0.0005
$\beta_{A7}$			-0.0293**	-0.0010*	-0.0005	-0.0052***
$\beta_{A8}$			-0.0178	-0.0003	-0.0027***	-0.0046**
$\beta_{A9}$			-0.0114	-0.0011***	-0.0031***	-0.0042
$\beta_{A10}$			-0.0095	-0.001***	-0.0023***	-0.0052**
$\beta_1$	0.0861	-0.0265	-0.0093***	0.0027***	0.0006*	0.0004***
$\beta_2$	-5.8171	1.6171	0.4509**	-0.0748***	-0.0099***	-0.0012
$R_a^2$	4.89%	3.04%	5.32%	10.20%	21.36%	7.93%

are driven by uncertainty about speculators' endowments in the on-the-run securities (Corollary 1).

According to our theory, government auctions are the critical events discriminating among those otherwise identical assets. In the stylized economy of Section 2.1, liquidity in both markets (1 and 2) is driven by MMs' perceived adverse selection risk in the presence of speculators trading strategically with their private fundamental information. However, the same speculators face uninformative, undisclosed endowment shocks in only one of the assets ( $e_2(k)$ ). We interpret these shocks as government auction allocations, and that asset (asset 2) as the just-issued—hence by definition on-the-run—security. Speculators are assumed to care about both their short- and long-term wealth, hence they care about the interim value of these allocations as well. Therefore, their trading activity in the on-the-run security also depends upon those uninformative endowment shocks, i.e., is informationally suboptimal. This attenuates MMs' adverse selection in that market, the more so the more short-term wealth matters to speculators ( $\gamma$ ) and the greater is the uncertainty surrounding their endowments ( $\sigma_e^2$ ), ultimately improving that market's liquidity with respect to the market for the off-the-run asset (asset 1).

These results, summarized in Corollary 1, translate naturally into a testable conjecture in fixed income markets. This conjecture stems from the observation that uncertainty about speculators' endowments ( $\sigma_e^2$ ) is likely to be the greatest—hence the on-the-run liquidity phenomenon the most intense—at the completion of an auction and *declining* afterward, i.e., when market participants can learn from observed price movements about those endowments. We test for this possibility by estimating, for every bill and note in our sample, the

following amended specification of Eq. (10):

$$\Delta S_t = \beta_0 + \sum_{i=1}^N \beta_{Ai} \text{Auction}_{t-i} + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \varepsilon_t, \quad (11)$$

where  $\text{Auction}_{t-i}$  is a dummy variable equal to one on day  $t$  if day  $t - i$  is the most recent auction date for the corresponding bond and equal to zero otherwise. We choose  $N = 4$  for three-month and six-month bills and  $N = 10$  for all other bonds to prevent each post-auction window to overlap with the window of subsequent auction dates.<sup>30</sup> We cannot estimate the contemporaneous impact of the auction on  $\Delta S_t$  ( $i = 0$ ) since GovPX reports transaction data on the auctioned on-the-run and the just off-the-run bonds only from the first business day *after* the auction date onward. Hence, we interpret estimates of  $\beta_0$ —the mean spread differential over the unaccounted portion of the prior auction cycle ending *on* day  $t$ —as a proxy for the extent of the on-the-run phenomenon immediately before trading on the new bond begins. Those estimates, reported in Table 4, indicate that, consistent with our conjecture (and Corollary 1), the off/on-the-run bid–ask spread differential  $\Delta S_t$  is *smaller* immediately following on-the-run auction dates: estimated coefficients  $\beta_{Ai}$  in Eq. (11) are negative and significant for both bills and notes, albeit often first

<sup>30</sup> The minimum number of days between auctions is in fact four days for three-month and six-month bills and between 14 and 22 days for all other bills and notes. Yet, similar inference ensues from either smaller or bigger but non-overlapping post-event windows. To control for auction reopenings, which are rare in our sample, we augment Eq. (11) with a dummy variable equal to one on auction reopening days and zero otherwise. See Fabozzi and Fleming (2004) for a detailed description of the functioning of U.S. Treasury auctions.

**Table 5**

Further determinants of the off/on-the-run bid–ask spread differential.

This table reports ordinary least squares (OLS) estimates of the coefficient  $\beta_3$  from the following regression model (Eq. (12)) over the full sample (1/1992–12/2000):

$$\Delta S_t = \beta_0 + \sum_{i=1}^N \beta_{Ai} \text{Auction}_{t-i} + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \beta_3 X_t + \varepsilon_t,$$

where  $\Delta S_t = S_t^{\text{off}} - S_t^{\text{on}}$  is the daily average off/on-the-run percentage bid–ask discount (for U.S. Treasury bills) or price (for U.S. Treasury notes) spread differential multiplied by 100,  $\Delta D_t = D_t^{\text{off}} - D_t^{\text{on}}$  is the off/on-the-run modified duration differential,  $\Delta C_t = C_t^{\text{off}} - C_t^{\text{on}}$  is the off/on-the-run convexity differential,  $\text{Auction}_{t-i}$  is a dummy variable equal to one on day  $t$  if day  $t - i$  is the most recent auction date for the corresponding bond and equal to zero otherwise, and  $X_t = HL_t$ , the competitive yield range,  $Ten_t$ , the total amount tendered in the corresponding auction,  $Acc_t$ , the total amount accepted at the auction, or  $Vol_t$ , the Eurodollar implied volatility. We assume  $N = 4$  for three-month and six-month bills and  $N = 10$  for all other bonds.  $R_a^2$  is the adjusted  $R^2$  from the estimation of the fully specified regression above. A \*, \*\*, or \*\*\* indicates statistical significance at the 10%, 5%, or 1% levels, respectively, using Newey–West standard errors.

	Three-month	Six-month	One-year	Two-year	Five-year	10-year
$X_t$	$\beta_3$	$\beta_3$	$\beta_3$	$\beta_3$	$\beta_3$	$\beta_3$
$HL_t$	3.0594***	12.545***	12.109***	0.0227	−0.0057	−0.0634**
$R_a^2$	6.61%	5.41%	10.02%	10.45%	22.01%	8.50%
$Ten_t$	−0.0050***	−0.0012**	−0.0027***	−0.0001***	−0.0003***	−0.0002**
$R_a^2$	12.07%	3.64%	9.66%	14.05%	24.33%	8.33%
$Acc_t$	−0.0201***	−0.0079**	−0.0162***	−0.0004***	0.0005**	0.0005**
$R_a^2$	7.70%	3.45%	9.62%	15.27%	23.55%	8.32%
$Vol_t$	0.0041	0.0043**	0.0024**	0.0002***	0.0003***	0.0003***
$R_a^2$	5.44%	3.99%	5.66%	16.34%	30.69%	8.86%
$HL_t$	−1.2610	10.232**	9.7235***	−0.0284	−0.0159**	0.0387
$Ten_t$	−0.0062***	−0.0007	0.0002	−0.0002***	−0.0003***	−0.0001
$Acc_t$	−0.0147***	−0.0092**	−0.0110***	−0.0002**	0.0010***	0.0008***
$Vol_t$	0.0087***	0.0059***	0.0032***	0.0003***	0.0004***	0.0003***
$R_a^2$	16.07%	7.13%	12.33%	27.87%	37.74%	9.49%

increasing and then decreasing in absolute magnitude. Hence, average liquidity differentials  $\beta_0 + \beta_{Ai}$  are generally lower in the immediate aftermath of Treasury auctions, albeit often less so thereafter.<sup>31</sup>

Perhaps a better proxy for endowment uncertainty induced by Treasury auctions is the range of competitive yield bids  $HL_t$ , defined in Section 3.1 as the ratio of the difference between the highest and lowest bid at an auction and the average accepted competitive bid.<sup>32</sup> As such, this variable measures both the variance in the demand for the auctioned security and, assuming market clearance, the variance in government debt endowed. Thus, ceteris paribus, the greater is the ratio  $HL_t$  the greater is the uncertainty among uninformed market participants about the final outcome of the auction for each of the sophisticated speculators, the greater is the uncertainty about their endowments of on-the-run bonds ( $\sigma_e^2$ ), hence the greater is the resulting off/on-the-run

liquidity differential. We test for this possibility by amending Eq. (11) as follows:

$$\Delta S_t = \beta_0 + \sum_{i=1}^N \beta_{Ai} \text{Auction}_{t-i} + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \beta_3 X_t + \varepsilon_t, \quad (12)$$

where  $X_t = HL_t$ . We report estimates of  $\beta_3$  in Table 5 for each of the securities in our sample. Estimates of all other coefficients are both qualitatively and quantitatively similar to those in Table 4, hence are not reported here. Table 5 shows that, consistent with Corollary 1, the competitive yield range  $HL_t$  is strongly positively related to the liquidity differential of Treasury bills—even after controlling for supply effects and fundamental volatility (see the discussion next)—yet is instead mostly unrelated to the liquidity differential of Treasury notes. To interpret this mixed evidence, we observe that the time when assets 1 and 2's identical payoffs  $v$  are realized in our model ( $t = 2$ ) can be thought of as the time when two identical bonds mature. Ceteris paribus, it is then reasonable to conjecture that the distinction between short- and long-term should be more relevant for Treasury notes than for bills, i.e., that  $\gamma \approx 1 - \gamma$  for speculators in the latter. Accordingly, Table 5 suggests that the effect of

<sup>31</sup> Consistently, Goldreich, Hanke and Nath (2005) show that average daily quoted and effective bid–ask spreads over the first 100 trading days of newly issued two-year Treasury notes (Fig. 2A) are first declining, then flat, and eventually steadily widening afterward.

<sup>32</sup> This information is announced by the U.S. Treasury at around 1 p.m. on the auction date.

uncertainty about speculators' endowments of the on-the-run asset ( $e_2(k)$ ) on  $\Delta S_t$  is greater (and in the direction of the theory) for Treasury securities of shorter maturity.<sup>33</sup>

The relative supply of new and old Treasury securities in the secondary market, as well as the demand for the new ones in the primary market, do not play any role in the stylized model of trading à la Kyle (1985) of Section 2. Nonetheless, these effects may intuitively contribute to the dynamics of the bid–ask spread differentials reported in Table 4. For instance, according to Vayanos and Weill (2008), the ensuing search costs—such as the additional time it may take a trader to locate a scarce off-the-run issue over its abundant on-the-run bond—may translate into liquidity wedges and no-arbitrage price premiums. We assess the relevance of these considerations by amending the above regression to include either the total amount tendered at the Treasury auctions ( $X_t = Ten_t$ ), the total amount eventually accepted by the investors ( $X_t = Acc_t$ ), or both.<sup>34</sup> The resulting estimated parameters, not reported here, indicate that our inference is robust to the inclusion of supply and demand effects: sign, magnitude, and significance of the coefficients  $\beta_{oi}$  are very similar to those displayed in Table 4. Consistent with the intuition above, estimates of  $\beta_3$  (in Table 5) are in most cases negative and significant: Tendered and accepted amounts lower bid–ask spread differentials in the Treasury market. Yet, their inclusion improves only marginally the overall fit (i.e., the adjusted  $R^2$ ,  $R_a^2$ ) of the regressions in Table 4.

### 4.3. The informational role of trading

The evidence reported in Section 4.2 provides further, more direct support for the basic premise of our model, i.e., that uncertainty surrounding speculators' endowments of new, just-auctioned securities creates a liquidity wedge between those securities and otherwise identical, old securities. Given this crucial premise, we now test two additional predictions of our theory that stem from the informational role of trading in our stylized model. These predictions are unique to that model, i.e., cannot easily be attributed to the alternative explanations for the on-the-run liquidity phenomenon discussed in Section 4.1. As such, if validated, they provide further, indirect support for our theory.

The first one (again from Corollary 1) states that, ceteris paribus, greater uncertainty surrounding both on-the-run and off-the-run assets' terminal payoffs (higher  $\sigma_v^2$ ) leads to greater liquidity differentials between them, for adverse selection risk becomes more severe for uninformed MMs in both assets, yet the more so in the

off-the-run security (asset 1) where noise trading is less intense ( $\sigma_z^2 < \sigma_n^2$ ). To evaluate this argument, we amend Eq. (12) by imposing that  $X_t = Vol_t$ , the daily Eurodollar implied volatility from Bloomberg, a commonly used proxy for the market's perceived uncertainty surrounding U.S. monetary policy. We report estimates of the corresponding coefficients  $\beta_3$  in Table 5. Consistent with Corollary 1 and the discussion in the previous section, greater Eurodollar implied volatility translates into greater off/on-the-run liquidity differentials: estimated  $\beta_3$  are always positive, always statistically significant at the 5% level or better (with the exception of three-month bills), and (relatively) larger for bills than for notes—i.e., when  $\gamma \approx 1 - \gamma$  (see Section 4.2) and so  $\sigma_n^2 > \sigma_z^2$  (see Section 2.1.3). These coefficients are even larger after controlling for supply effects and endowment uncertainty, in the bottom panel of Table 5.

The second prediction (from Remark 1) states that because of the informational role of trading in the markets for asset 1 and asset 2, the degree of heterogeneity of speculators' private information has an impact on the equilibrium liquidity differential between those markets whose sign depends on speculators' relative numerosity ( $M$  in Eq. (7)). We test for this argument by amending Eq. (11) as follows:

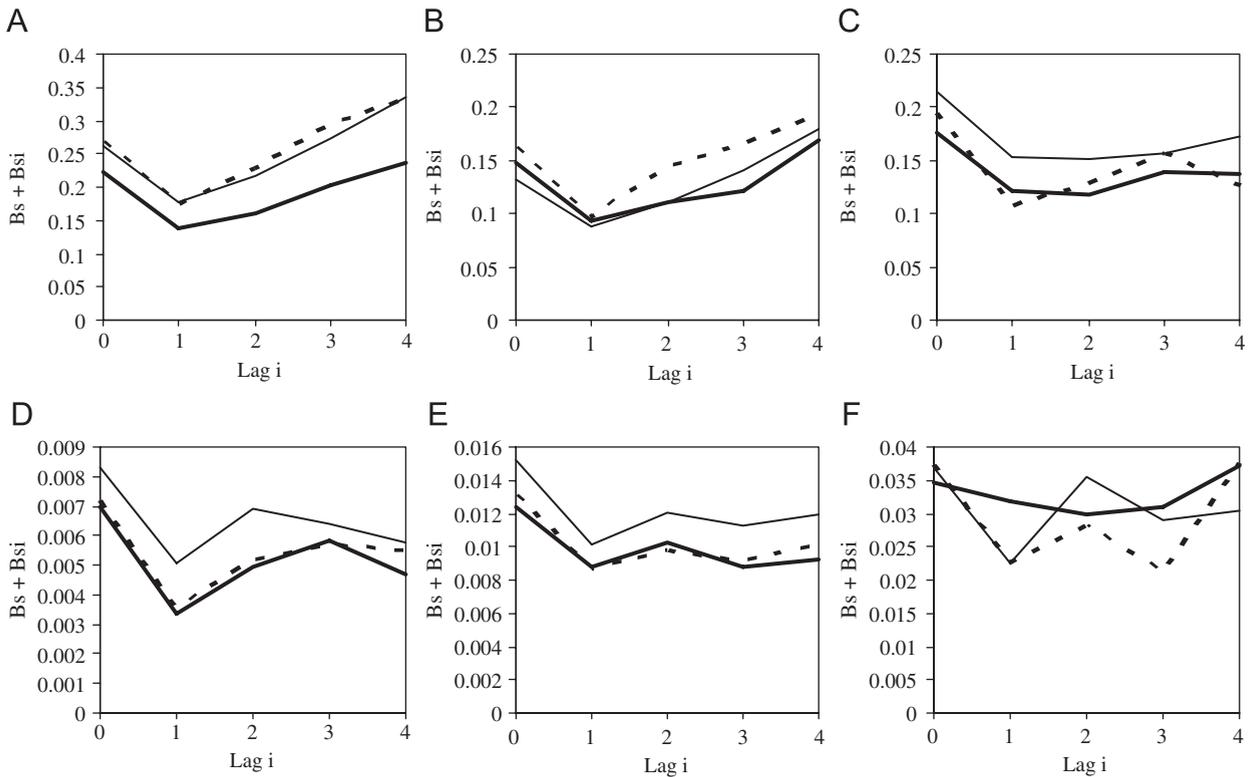
$$\begin{aligned} \Delta S_t = & \beta_h D_{ht} + \beta_l D_{lt} + \beta_m (1 - D_{ht} - D_{lt}) \\ & + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \sum_{i=1}^N \beta_{hi} D_{ht} Auction_{t-i} \\ & + \sum_{i=1}^N \beta_{li} D_{lt} Auction_{t-i} \\ & + \sum_{i=1}^N \beta_{mi} (1 - D_{ht} - D_{lt}) Auction_{t-i} + \varepsilon_t, \end{aligned} \quad (13)$$

where  $D_{ht}$  ( $D_{lt}$ ) is a dummy variable equal to one on days with high (low) information heterogeneity, defined in Section 3.2 as days in which the monthly variable  $SSD_{Pt}$  of Eq. (9) is above (below) the top (bottom) 70th (30th) percentile of its empirical distribution, and equal to zero otherwise. We compute  $SSD_{Pt}$  using all the announcements listed in Table 2 (i.e.,  $P = 18$  in Eq. (9)). We obtain qualitatively similar results for  $P = 1$  (Nonfarm Payroll) or  $P = 7$  (the influential announcements listed in Section 3.2), as well as by including security-pair and calendar fixed effects in Eq. (9) and/or repo rate differentials (as in Table 3).

For conciseness' sake, we only show plots of the resulting estimated average liquidity differentials  $\beta_h + \beta_{hi}$ ,  $\beta_m + \beta_{mi}$ , and  $\beta_l + \beta_{li}$  for  $i = 1, 2, 3$ , and 4 and for each of the bills and notes in our sample in Fig. 3. As already suggested by Table 4, off/on-the-run bid–ask spread wedges are lower right after Treasury auction dates regardless of the degree of information heterogeneity among speculators, again consistent with Corollary 1. Fig. 3 also suggests that those liquidity differentials are sensitive to the degree of information heterogeneity about macroeconomic fundamentals among sophisticated market participants, consistent with Remark 1. In particular, average  $\Delta S_t$  is generally increasing (i.e.,  $\beta_{hi} > \beta_{li}$ ) in the heterogeneity of speculators' beliefs (i.e., decreasing in  $\rho$  in Fig. 1A), often statistically

<sup>33</sup> In unreported analysis, the same inference can be drawn when accounting for the interaction of  $HL_t$  with the auction dummies  $Auction_{t-i}$  in Eq. (12), as well as when controlling for repo rate differentials and/or year dummies. However, not surprisingly, allowing for security-pair fixed effects weakens the statistical significance of any auction characteristic (including  $HL_t$ ) in Eq. (12).

<sup>34</sup> Unreported analysis indicates that similar inference ensues from the inclusion of the interaction of both variables with auction dummies  $Auction_{t-i}$ , in Eq. (12), as well as from the inclusion of security-pair fixed effects, year dummies, and/or repo rate differentials (as in Table 3).



**Fig. 3.** Information heterogeneity and the off/on-the-run bid-ask spread differential. In this figure, we plot sums of coefficients  $\beta_h + \beta_{hi} \times 100$  (thin line),  $\beta_m + \beta_{mi} \times 100$  (solid line), and  $\beta_l + \beta_{li} \times 100$  (dotted line) from the ordinary least squares (OLS) estimation of the following regression model (Eq. (13)) over the full sample (1/1992–12/2000) for three-month (A), six-month (B), and one-year (C) U.S. Treasury bills, and two-year (D), five-year (E), and 10-year (F) U.S. Treasury notes.

$$\Delta S_t = \beta_h D_{ht} + \beta_l D_{lt} + \beta_m (1 - D_{ht} - D_{lt}) + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \sum_{i=1}^N \beta_{hi} D_{ht} \text{Auction}_{t-i} + \sum_{i=1}^N \beta_{li} D_{lt} \text{Auction}_{t-i} + \sum_{i=1}^N \beta_{mi} (1 - D_{ht} - D_{lt}) \text{Auction}_{t-i} + \varepsilon_t,$$

where  $\Delta S_t = S_t^{\text{off}} - S_t^{\text{on}}$  is the daily average off/on-the-run percentage bid-ask discount (for U.S. Treasury bills) or price (for U.S. Treasury notes) spread differential multiplied by 100,  $\Delta D_t = D_t^{\text{off}} - D_t^{\text{on}}$  is the off/on-the-run modified duration differential,  $\Delta C_t = C_t^{\text{off}} - C_t^{\text{on}}$  is the off/on-the-run convexity differential,  $\text{Auction}_{t-i}$  is a dummy variable equal to one on day  $t$  if day  $t - i$  is the most recent auction date for the corresponding bond and equal to zero otherwise,  $D_{ht}$  ( $D_{lt}$ ) is a dummy variable equal to one on days with high (low) information heterogeneity, defined in Section 3.2 as days in which the monthly variable  $SSD_{pt}$  of Eq. (9) is above (below) the top (bottom) 70th (30th) percentile of its empirical distribution, and equal to zero otherwise, and  $P = 18$ , for each of the bills and note in our sample. We assume  $N = 4$  for three-month and six-month bills and  $N = 10$  for all other bonds.

significantly so, for issues of longer maturity (one-year bills to 10-year notes). This is intuitive since, according to our model, more heterogeneously informed speculators trade more cautiously to protect their perceived private information monopoly, the more so in the less liquid market (off-the-run), thus widening its liquidity gap with the on-the-run market. Yet, average spread differentials are either insensitive to or even weakly increasing in  $\rho$  (i.e.,  $\beta_{hi} \lesssim \beta_{li}$ ) for short-term bills. According to our model (see Fig. 1A), this dichotomy may be explained by Treasury bills' markets being populated by fewer, hence less competitive sophisticated speculators. Anecdotal evidence, the significantly wider bid-ask spreads and lower aggregate daily trading volume and trading frequency in bills than in notes (e.g., our Tables 1A and 1B, and Fleming, 2003, Tables 1 and 2), as well as the observation that informed investors may be more active in more liquid trading venues (e.g., Chowdhry and Nanda, 1991) suggest that this may indeed be the case.

Overall, the above results provide additional support for our model, for they indicate that the magnitude and

dynamics of off/on-the-run liquidity differentials—which we showed to be related to endowment uncertainty following on-the-run auctions in Section 4.2—are also crucially related to the informational role of trading in the U.S. Treasury market.

#### 4.4. Macroeconomic news

Macroeconomic news is frequently released to the public in the U.S. financial markets. For instance, more than 2,000 of the news items listed in Table 2 were announced, often on the same day, over our sample period. These news releases are especially relevant for the U.S. Treasury market since their potential information content is deemed to play a crucial role for the valuation of government bonds. Consistently, Pasquariello and Vega (2007) find that the release of macroeconomic information (weakly) improves liquidity in the Treasury note market. According to our model, these news releases

**Table 6**

Macroeconomic news and the on-the-run liquidity phenomenon.

This table reports ordinary least squares (OLS) estimates of the differences  $(\beta_{0w}^{ann} - \beta_{0w}^{noann}) \times 100$  from the following regression model (Eq. (14)) over the full sample (1/1992–12/2000):

$$\Delta S_t = \sum_{w=1}^5 \beta_{0w}^{ann} d_{tw} Ann_t + \sum_{d=1}^5 \beta_{0w}^{noann} d_{tw}(1 - Ann_t) + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \varepsilon_t,$$

where  $Ann_t$  is a dummy variable equal to one if any of the seven influential announcements listed in Section 3.2 ( $P = 7$ ) is released on day  $t$  and equal to zero otherwise,  $d_{tw}$  are day-of-week dummy variables from  $w = 1$  (Monday) to 5 (Friday),  $\Delta S_t = S_t^{off} - S_t^{on}$  is the daily average off/on-the-run percentage bid-ask discount (for U.S. Treasury bills) or price (for U.S. Treasury notes) spread differential multiplied by 100,  $\Delta D_t = D_t^{off} - D_t^{on}$  is the off/on-the-run modified duration differential, and  $\Delta C_t = C_t^{off} - C_t^{on}$  is the off/on-the-run convexity differential. A \*, \*\*, or \*\*\* indicates statistical significance of the  $F$ -statistic for the corresponding difference at the 10%, 5%, or 1% levels, respectively, using Newey-West standard errors.

	$(\beta_{0w}^{ann} - \beta_{0w}^{noann})$				
	Monday	Tuesday	Wednesday	Thursday	Friday
Three-month	2.589	-3.357*	1.158	8.652***	1.277
Six-month	0.843	-0.784	1.401	6.570***	5.233***
One-year	0.876	-0.468	-0.444	5.705***	-0.024
Two-year	-0.116***	-0.002	-0.013	0.066	0.054
Five-year	0.016	0.044	0.093*	0.106	0.070
10-year	0.111	0.042	0.086	0.478**	0.181

may be relevant for the on-the-run liquidity phenomenon as well. In particular, we show in Section 2.2 that the availability of a public signal of the identical terminal payoff of both the off-the-run and the on-the-run securities ( $v$ ) reduces their liquidity differentials—the more so the better is the quality of that signal—for it attenuates both markets' adverse selection risk, yet mainly where most severe (the off-the-run market).

We assess the empirical relevance of these considerations by using the database of macroeconomic announcements described in Section 3.2. Specifically, the above implications translate into observing a negative difference between each  $\beta_{0w}^{ann}$  and  $\beta_{0w}^{noann}$  in the following amended specification of Eq. (10):

$$\Delta S_t = \sum_{w=1}^5 \beta_{0w}^{ann} d_{tw} Ann_t + \sum_{w=1}^5 \beta_{0w}^{noann} d_{tw}(1 - Ann_t) + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \varepsilon_t, \tag{14}$$

where  $Ann_t$  is a dummy variable equal to one if either the Nonfarm Payroll Employment report ( $P = 1$ ), any of the seven influential announcements listed in Section 3.2 ( $P = 7$ ), or any of the 18 announcements listed in Table 2 ( $P = 18$ ) is released on day  $t$  and equal to zero otherwise, while  $d_{tw}$  are day-of-week dummy variables, from  $w = 1$  (Monday) to 5 (Friday), to control for event-day clustering. We report the resulting estimates of day-specific differences in Table 6 for each of the bonds in our sample when  $P = 7$ .<sup>35</sup> We discuss the estimates for  $P = 1$  or 18 below.<sup>36</sup>

At first sight, the evidence in Table 6 is unsupportive of Corollary 2. Estimates of  $\beta_{0w}^{ann} - \beta_{0w}^{noann}$  are in fact negative much less frequently than positive and most often

statistically indistinguishable from zero. This can be due to several factors. Extant theories suggest alternative mechanisms mitigating the impact of the availability of public signals on  $\Delta S_t$ . For instance, according to Chowdhry and Nanda (1991) sophisticated investors may divert much of their trading activity to the most liquid venue to maximize their expected profits. In such a setting, the release of high-quality public information, by devaluing those investors' private signals, may make that migration even more intense, thus *widening*—rather than *tightening*, as instead argued in Section 2.2.1—the equilibrium liquidity differentials among markets. In addition, both the dispersion of beliefs among market participants and the quality of available public signals might vary across announcements, ultimately influencing the net effect of their arrival on  $\Delta S_t$ . For example, Kim and Verrecchia (1994) argue that, in the presence of endogenous information acquisition, market liquidity may *deteriorate* when public signals are released, but unequivocally improves with greater precision of those signals and less private information heterogeneity. According to our model (Corollary 2 and Remark 2), both factors affect sign and significance of the relation between the availability of public signals and the on-the-run liquidity phenomenon. The weaker statistical significance of estimates for  $\beta_{0w}^{ann} - \beta_{0w}^{noann}$  for the narrowest and broadest—hence of possibly the highest and lowest quality—sets of macroeconomic news (i.e., for  $P = 1$  and 18, not reported here) provides preliminary support to both sets of arguments above, respectively.

To test for the relevance of these considerations, we proceed in two directions. First, we focus on the impact of public signal noise ( $\sigma_p^2$  of Section 2.2) on  $\Delta \lambda^* - \Delta \lambda$  of Eq. (8). We measure  $\sigma_p^2$  using the U.S. government's frequent revisions of previously released macroeconomic information, as in Pasquariello and Vega (2007) and Aruoba (2008). Specifically, we augment our database with the Federal Reserve Bank of Philadelphia "Real Time Data Set" (RTDS) of all "informative" monthly data

<sup>35</sup> As a word of caution, we observe that one of the seven influential news in the MMS database, the Initial Unemployment Claims report, is released weekly in all but 24 Thursdays in our sample. Hence, when  $P = 7$  or 18, the coefficient  $\beta_{04}^{noann}$  is estimated with only 24 observations.

<sup>36</sup> Unreported analysis shows the inference discussed in this section to be unaffected by the inclusion of security-pair and calendar fixed effects in Eq. (14) and/or repo rate differentials, as discussed in Section 4.

**Table 7**

Public signal noise and the on-the-run liquidity phenomenon.

This table reports ordinary least squares (OLS) estimates of the differences ( $\beta_{0w}^{ann} - \beta_{0w}^{noann}$ ) from the following regression model (Eq. (14)) over the full sample (1/1992–12/2000):

$$\Delta S_t = \sum_{w=1}^5 \beta_{0w}^{ann} d_{tw} Ann_t + \sum_{d=1}^5 \beta_{0d}^{noann} d_{tw}(1 - Ann_t) + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \varepsilon_t,$$

where  $Ann_t$  is a dummy variable equal to one if Industrial Production news is released on day  $t$  and equal to zero otherwise,  $d_{tw}$  are day-of-week dummy variables, from  $w = 1$  (Monday) to 5 (Friday),  $\Delta S_t = S_t^{off} - S_t^{on}$  is the daily average off/on-the-run percentage bid–ask discount (for U.S. Treasury bills) or price (for U.S. Treasury notes) spread differential multiplied by 100,  $\Delta D_t = D_t^{off} - D_t^{on}$  is the off/on-the-run modified duration differential, and  $\Delta C_t = C_t^{off} - C_t^{on}$  is the off/on-the-run convexity differential. Eq. (14) is estimated separately in days in which the noise surrounding the announcements listed above is high (Panel A) and low (Panel B) for each of the bills and notes in our sample. Specifically, we measure public signal noise as the absolute difference between each initial announcement and its last revision. We then label the corresponding announcement days as characterized by high (low) noise when that difference is in the top (bottom) 70th (30th) percentile of its empirical distribution. A \*, \*\*, or \*\*\* indicates statistical significance of the  $F$ -statistic for the corresponding difference at the 10%, 5%, or 1% levels, respectively, using Newey–West standard errors.

	Panel A: ( $\beta_{0w}^{ann} - \beta_{0w}^{noann}$ ) when $\sigma_p^2$ is high					Panel B: ( $\beta_{0w}^{ann} - \beta_{0w}^{noann}$ ) when $\sigma_p^2$ is low				
	Monday	Tuesday	Wednesday	Thursday	Friday	Monday	Tuesday	Wednesday	Thursday	Friday
Three-month	27.815***	-3.043	-7.649***	-5.276	-3.811	-16.532***	-1.546	2.225	-20.171***	3.786
Six-month	-10.239***	-2.192	-5.855***	-6.201***	-5.387	-10.552***	1.982	-1.144	-4.905***	3.369
One-year	12.470***	10.802	8.390	7.178**	8.518*	-6.378***	-1.793	0.853	-10.333***	5.423
Two-year	0.129	-0.127	0.385***	-0.255***	-0.005	0.010	0.150	-0.112	-0.097***	-0.072
Five-year	0.209	0.095	0.343	-0.078	-0.007	0.036	-0.030	-0.290*	-0.190***	0.063
10-year	-1.022**	0.307	1.283	0.204	-0.360	-0.459	-0.021	0.417	-0.757***	0.422

revisions (i.e., those due to newly available information).<sup>37</sup> These revisions are available to us only for Capacity Utilization, Industrial Production, and Nonfarm Payroll Employment, among the 18 news releases listed in Table 2. We then compute those public signals' noise as the absolute difference between each initial announcement and its last revision and label the corresponding announcement days as characterized by high (low) noise  $\sigma_p^2$  when that difference is in the top (bottom) 70th (30th) percentile of its empirical distribution.<sup>38</sup> Lastly, we estimate Eq. (14) for each of the RTDS announcements in either of their corresponding subsets of high and low  $\sigma_p^2$  days in our sample. We report the resulting differences  $\beta_{0w}^{ann} - \beta_{0w}^{noann}$  for Industrial Production in Table 7.

These estimates are striking: average  $\Delta S_t$  during Industrial Production announcement days is lower than during non-announcement days more often, more so, and more significantly so when the quality of that announcement is higher. Specifically, the statistically significant differences  $\beta_{0w}^{ann} - \beta_{0w}^{noann}$  in Panel B (low  $\sigma_p^2$  announcement days) are always negative, while those in Panel A (high  $\sigma_p^2$  announcement days) are often positive, especially in the

market for Treasury bills—i.e., where we conjectured the distinction between short- and long-term to be least relevant ( $\gamma \approx 1 - \gamma$ ), hence the underlying adverse selection differential between on-the-run and off-the-run securities most severe (see Section 4.2). The inference drawn upon Capacity Utilization announcement days (not reported here) is qualitatively and quantitatively similar. However, we did not find any meaningful differences in  $\beta_{0w}^{ann} - \beta_{0w}^{noann}$  when estimated in correspondence with Nonfarm Payroll announcement days (also not reported here). This is not surprising, in light of the potentially offsetting liquidity-migration effect discussed in Chowdhry and Nanda (1991), since those news releases are commonly characterized as of the highest and most homogeneous quality.<sup>39</sup> Thus, Table 7 suggests that the decline in off/on-the-run bid–ask spread differentials in the presence of a public signal is both more economically and statistically significant when  $\sigma_p^2$  is low than when  $\sigma_p^2$  is high, consistent with our theory.

Second, Remark 2 states (and Fig. 1B shows) that in the presence of a public signal of the traded assets' fundamentals ( $S_p$ ), the decline in the resulting off/on-the-run liquidity differential is the greatest when information heterogeneity is the highest ( $\rho$  is lowest) among sophisticated investors in the venues when the latter are most numerous, i.e., in the Treasury notes' markets (as argued in Section 4.3). Intuitively, adverse selection is most severe in the off-the-run market (asset 1) when many speculators are most cautious (low  $\rho$ ), hence the benefit of  $S_p$ 's availability for the MMs is the greatest. We assess this argument by estimating Eq. (14) over the subset of days in our sample characterized by high (low) information

<sup>37</sup> Occasionally, the U.S. government performs “uninformative” revisions of its previously announced data, i.e., due to definitional changes (such as changes in the base-year or changes in seasonal weights). Over our sample period, Industrial Production was the only announcement undergoing one such “uninformative” change, a base-year revision in February 1998. For a more detailed description of the RTDS dataset and its properties, see Croushore and Stark (2001).

<sup>38</sup> By definition, the final published revision of an announcement represents the most accurate measure for the corresponding macro-economic variable. The above procedure is motivated by the observation that these revisions can be interpreted as noise since they are predictable based on past information (e.g., Mork, 1987; Faust, Rogers, and Wright, 2005; Aruoba, 2008). Pasquariello and Vega (2007) find a more pronounced improvement in Treasury notes' market liquidity when low noise announcements are released to the public.

<sup>39</sup> For example, Andersen and Bollerslev (1998), among others, label the Nonfarm Payroll report the “king” of announcements for its release has the most significant impact on most asset markets.

**Table 8**

Public signal, information heterogeneity, and the on-the-run liquidity phenomenon.

This table reports ordinary least squares (OLS) estimates of the differences  $(\beta_{0w}^{ann} - \beta_{0w}^{noann}) \times 100$  from the following regression model (Eq. (14)) over the full sample (1/1992–12/2000):

$$\Delta S_t = \sum_{w=1}^5 \beta_{0w}^{ann} d_{tw} Ann_t + \sum_{d=1}^5 \beta_{0w}^{noann} d_{tw} (1 - Ann_t) + \beta_1 \Delta D_t + \beta_2 \Delta C_t + \varepsilon_t,$$

where  $Ann_t$  is a dummy variable equal to one if either Capacity Utilization, Industrial Production, or Nonfarm Payroll Employment is released on day  $t$  and equal to zero otherwise,  $d_{tw}$  are day-of-week dummy variables, from  $w = 1$  (Monday) to 5 (Friday),  $\Delta S_t = S_t^{off} - S_t^{on}$  is the daily average off/on-the-run percentage bid–ask discount (for U.S. Treasury bills) or price (for U.S. Treasury notes) spread differential multiplied by 100,  $\Delta D_t = D_t^{off} - D_t^{on}$  is the off/on-the-run modified duration differential, and  $\Delta C_t = C_t^{off} - C_t^{on}$  is the off/on-the-run convexity differential. Eq. (14) is estimated separately in days in which information heterogeneity among speculators is high ( $\rho$  is low, in Panel A) and low ( $\rho$  is high, in Panel B) for each of the bills and notes in our sample. Specifically, we define high (low) information heterogeneity days as days in which the monthly variable  $SSD_{pt}$  of Eq. (9) is above (below) the top (bottom) 70th (30th) percentile of its empirical distribution when  $P = 7$  (see Section 3.2). A \*, \*\*, or \*\*\* indicates statistical significance of the  $F$ -statistic for the corresponding difference at the 10%, 5%, or 1% levels, respectively, using Newey–West standard errors.

	Panel A: $(\beta_{0w}^{ann} - \beta_{0w}^{noann})$ when $\rho$ is low					Panel B: $(\beta_{0w}^{ann} - \beta_{0w}^{noann})$ when $\rho$ is high				
	Monday	Tuesday	Wednesday	Thursday	Friday	Monday	Tuesday	Wednesday	Thursday	Friday
Three-month	–2.826	–2.423	3.974	2.041	–2.748	2.869	–5.990	1.307	–0.030	–5.410
Six-month	13.643	–1.528	1.023	2.674	–7.876	–2.224	–0.479	2.130	3.994	3.944
One-year	–6.519	–0.670	–0.521	–2.585	–4.158	2.154	–1.284	0.702	0.082	0.229
Two-year	–0.176**	0.082	–0.117	0.026	0.114	–0.070	0.046	0.026	0.058	0.023
Five-year	0.000	0.189*	0.064	0.005	0.216*	0.074	–0.015	0.210*	0.032	–0.200
10-year	–0.001	–0.057	–0.113	–0.255	0.070	0.326	0.473	0.097	–0.007	0.277

heterogeneity, defined in Section 3.2 as days in which the average dispersion of professional forecasts of  $P$  announcements from the MMS database— $SSD_{pt}$  of Eq. (9)—is above (below) the top (bottom) 70th (30th) percentile of its empirical distribution. We then report the ensuing differences  $\beta_{0w}^{ann} - \beta_{0w}^{noann}$  in Table 8 when  $P = 7$  and  $\rho$  is either low (Panel A) or high (Panel B). Consistent with Remark 2 and Fig. 1B, the estimated  $\beta_{0w}^{ann} - \beta_{0w}^{noann}$  are larger and more often negative—i.e., off/on-the-run bid–ask spread differentials  $\Delta S_t$  decline during announcement days—when speculators’ beliefs are more heterogeneous ( $SSD_{pt}$  is high, in Panel A), especially for longer-term bills and notes. Yet, since we are not cross-sorting announcement days by public signal noise (as in Table 7), most of these differences are again not statistically significant (as in Table 6). Qualitatively similar inference (not reported here) stems from  $P = 1$  or 18.

Overall, the above evidence indicates that, as postulated by our theory, the availability of public signals of assets’ terminal payoffs mitigates the on-the-run liquidity phenomenon in the U.S. Treasury market—which we model as and show to be related to auction-driven endowment uncertainty in Sections 2.1 and 4.2, respectively—by alleviating adverse selection among market participants.

### 5. Conclusions

The existence of a negative liquidity differential between on-the-run and off-the-run securities is a pervasive and not fully understood feature of both domestic and international fixed income markets. The main goal of this paper is to deepen our understanding of the links between this important aspect of the on-the-run phenomenon, news about fundamentals, and strategic

trading conditional on investors’ dispersion of beliefs and public signals’ noise.

To that end, we develop a parsimonious model of speculative trading in multiple assets in the presence of heterogeneously informed, imperfectly competitive traders, auction-driven endowment shocks identifying the on-the-run security from the off-the-run security, and a public signal of their identical terminal value. We then test its equilibrium implications by studying the determinants of daily differences in bid–ask spreads—a common and effective measure of bond market liquidity—for on-the-run and off-the-run three-month, six-month, and one-year U.S. Treasury bills and two-year, five-year, and 10-year U.S. Treasury notes.

Our evidence indicates that (i) the resulting off/on-the-run liquidity differentials are large, even after controlling for several differences in their intrinsic characteristics (such as duration, convexity, repo rates, or term premiums), and (ii) an economically meaningful portion of those liquidity differentials is linked to strategic trading activity in both security types. The nature of this linkage is sensitive to the uncertainty surrounding auction shocks and the economy, the intensity of investors’ dispersion of beliefs, and the noise of the public announcement. In particular, and consistent with our model, off/on-the-run liquidity differentials are smaller immediately following bond auction dates and in the presence of (high-quality) macroeconomic announcements, and larger when the dispersion of auction bids is higher, when fundamental uncertainty is greater, and when the beliefs of sophisticated traders are more heterogeneous.

These findings suggest that liquidity differentials between on-the-run and off-the-run securities depend crucially on endowment uncertainty in the former and the informational role of strategic trading in both. We believe this is an important implication for future

research (including our own) on the on-the-run price phenomenon.

## Appendix

**Proof of Proposition 1.** As noted in Section 2.1.2, the proof is by construction. We start by guessing that equilibrium  $p_{1,i}$  and  $x_i(k)$  are given by  $p_{1,i} = A_{0,i} + A_{1,i}\omega_{1,i}$  and  $x_i(k) = B_{0,i} + B_{1,i}\delta_v(k) + C_{1,i}e_2(k)$ , respectively, where  $A_{1,i} > 0$  and  $i = \{1, 2\}$ . Those expressions and the definition of  $\omega_{1,i}$  imply that, for the  $k$ th speculator,

$$E[p_{1,i}|\delta_v(k), e_2(k)] = A_{0,i} + A_{1,i}x_i(k) + A_{1,i}B_{0,i}(M-1) + A_{1,i}B_{1,i}(M-1)\rho\delta_v(k). \quad (\text{A.1})$$

Using Eq. (A.1), the first-order conditions of the maximization of the  $k$ th speculator's expected utility  $E_1^k[U(k)]$  with respect to  $x_1(k)$  and  $x_2(k)$  are given by

$$p_0 + \delta_v(k) - A_{0,1} - (M+1)A_{1,1}B_{0,1} - 2A_{1,1}B_{1,1}\delta_v(k) - (M-1)A_{1,1}B_{1,1}\rho\delta_v(k) - 2A_{1,1}C_{1,1}e_2(k) = 0, \quad (\text{A.2})$$

$$p_0 + \frac{\gamma}{1-\gamma}A_{1,2}e_2(k) + \delta_v(k) - A_{0,2} - (M+1)A_{1,2}B_{0,2} - 2A_{1,2}B_{1,2}\delta_v(k) - (M-1)A_{1,2}B_{1,2}\rho\delta_v(k) - 2A_{1,2}C_{1,2}e_2(k) = 0, \quad (\text{A.3})$$

respectively. The second-order conditions are satisfied, since  $2A_{1,i} > 0$ . For Eqs. (A.2) and (A.3) to be true, it must be that

$$p_0 - A_{0,1} = (M+1)A_{1,1}B_{0,1}, \quad (\text{A.4})$$

$$2A_{1,1}B_{1,1} = 1 - (M-1)A_{1,1}B_{1,1}\rho, \quad (\text{A.5})$$

$$2A_{1,1}C_{1,1} = 0, \quad (\text{A.6})$$

$$p_0 - A_{0,2} = (M+1)A_{1,2}B_{0,2}, \quad (\text{A.7})$$

$$2A_{1,2}B_{1,2} = 1 - (M-1)A_{1,2}B_{1,2}\rho, \quad (\text{A.8})$$

$$2A_{1,2}C_{1,2} = \frac{\gamma}{1-\gamma}A_{1,2}. \quad (\text{A.9})$$

Eqs. (A.6) and (A.9) imply that  $C_{1,1} = 0$  and  $C_{1,2} = \frac{1}{2}(\gamma/(1-\gamma))$ . The distributional assumptions of Section 2.1 imply that the order flows  $\omega_{1,1}$  and  $\omega_{1,2}$  are normally distributed with means  $E(\omega_{1,1}) = MB_{0,1}$  and  $E(\omega_{1,2}) = MB_{0,2}$ , and variances  $var(\omega_{1,1}) = MB_{1,1}^2\rho\sigma_v^2[1 + (M-1)\rho] + \sigma_z^2$  and  $var(\omega_{1,2}) = MB_{1,2}^2\rho\sigma_v^2[1 + (M-1)\rho] + \sigma_n^2$ , respectively. Since  $cov(v, \omega_{1,i}) = MB_{1,i}\rho\sigma_v^2$ , it ensues that

$$E(v|\omega_{1,1}) = p_0 + \frac{MB_{1,1}\rho\sigma_v^2}{MB_{1,1}^2\rho\sigma_v^2[1 + (M-1)\rho] + \sigma_z^2} \times (\omega_{1,1} - MB_{0,1}), \quad (\text{A.10})$$

$$E(v|\omega_{1,2}) = p_0 + \frac{MB_{1,2}\rho\sigma_v^2}{MB_{1,2}^2\rho\sigma_v^2[1 + (M-1)\rho] + \sigma_n^2} \times (\omega_{1,2} - MB_{0,2}). \quad (\text{A.11})$$

According to the definition of a Bayesian-Nash equilibrium in this economy (Section 2.1.1),  $p_{1,i} = E(v|\omega_{1,i})$ .

Therefore, our conjectures for  $p_{1,1}$  and  $p_{1,2}$  imply that

$$A_{0,1} = p_0 - MA_{1,1}B_{0,1}, \quad (\text{A.12})$$

$$A_{1,1} = \frac{MB_{1,1}\rho\sigma_v^2}{MB_{1,1}^2\rho\sigma_v^2[1 + (M-1)\rho] + \sigma_z^2}, \quad (\text{A.13})$$

$$A_{0,2} = p_0 - MA_{1,2}B_{0,2}, \quad (\text{A.14})$$

$$A_{1,2} = \frac{MB_{1,2}\rho\sigma_v^2}{MB_{1,2}^2\rho\sigma_v^2[1 + (M-1)\rho] + \sigma_n^2}. \quad (\text{A.15})$$

The expressions for  $A_{0,i}$ ,  $A_{1,i}$ ,  $B_{0,i}$ , and  $B_{1,i}$  in Proposition 1 must solve the system made of Eqs. (A.4), (A.5), (A.7), (A.8), and (A.12) to (A.15) to represent a linear equilibrium. Defining  $A_{1,1}B_{0,1}$  from Eq. (A.4) and  $A_{1,2}B_{0,2}$  from Eq. (A.7), and plugging them into Eqs. (A.12) and (A.14), respectively, leads us to  $A_{0,1} = A_{0,2} = p_0$ . Thus, it must be that  $B_{0,1} = B_{0,2} = 0$  to satisfy Eqs. (A.4) and (A.7). We are left with the task of finding  $A_{1,i}$  and  $B_{1,i}$ . Solving Eqs. (A.5) and (A.8) for  $A_{1,1}$  and  $A_{1,2}$ , respectively, we get

$$A_{1,1} = \frac{1}{B_{1,1}[2 + (M-1)\rho]}, \quad (\text{A.16})$$

$$A_{1,2} = \frac{1}{B_{1,2}[2 + (M-1)\rho]}. \quad (\text{A.17})$$

Equating Eqs. (A.16) and (A.17) to (A.13) and (A.15), respectively, it follows that  $B_{1,1}^2 = \sigma_z^2/M\rho\sigma_v^2$  and  $B_{1,2}^2 = \sigma_n^2/M\rho\sigma_v^2$ , i.e., that  $B_{1,1} = \sigma_z/\sqrt{M\rho}\sigma_v$  and  $B_{1,2} = \sigma_n/\sqrt{M\rho}\sigma_v$ . Substituting these expressions back into Eqs. (A.16) and (A.17) implies that  $A_{1,1} = \sqrt{M\rho}\sigma_v/\sigma_z[2 + (M-1)\rho]$  and  $A_{1,2} = \sqrt{M\rho}\sigma_v/\sigma_n[2 + (M-1)\rho]$ . Finally, we observe that Proposition 1 is equivalent to a symmetric Cournot equilibrium with  $M$  speculators. Therefore, the “backward reaction mapping” introduced by Novshek (1984) to find  $n$ -firm Cournot equilibria proves that, given any linear pricing rule, the symmetric linear strategies  $x_i(k)$  of Eqs. (5) and (6) indeed represent the unique Bayesian-Nash equilibrium of the Bayesian game among speculators.  $\square$

**Proof of Corollary 1.** The off/on-the run liquidity differential  $\Delta\lambda$  of Eq. (7) is positive since  $\sigma_z^2 < \sigma_n^2$  for any  $\gamma > 0$ . Furthermore,  $\partial\Delta\lambda/\partial\gamma = (\sqrt{M\rho}\sigma_v\sigma_z^2/\sigma_n^2\sigma_z^2[2 + (M-1)\rho]^2)M\sigma_e^2\gamma/4\sigma_n^2(1-\gamma)^3 > 0$ ,  $\partial\Delta\lambda/\partial\sigma_e = (\sqrt{M\rho}\sigma_v\sigma_z^2/\sigma_n^2\sigma_z^2[2 + (M-1)\rho]^2)M\gamma^2\sigma_e/4\sigma_n(1-\gamma)^2 > 0$ , and  $\partial\Delta\lambda/\partial\sigma_v = \sqrt{M\rho}(\sigma_n - \sigma_z)/\sigma_n\sigma_z[2 + (M-1)\rho] > 0$ .  $\square$

**Proof of Remark 1.** The statement stems from the fact that  $\partial\Delta\lambda/\partial\rho = \sigma_v\sigma_n(\sigma_n - \sigma_z)[2 - (M-1)\rho]/2\sqrt{\rho}\sigma_n^2\sigma_z^2[2 + (M-1)\rho]^2 \geq 0$  when  $\rho \leq 2/(M-1)$  and is negative otherwise. When  $M = 2$  or  $3$ ,  $\partial\Delta\lambda/\partial\rho$  is always positive since  $\rho \in (0, 1]$ . Yet, the greater is  $M$  the smaller is the subset of  $\rho \in (0, 1]$  such that  $\partial\Delta\lambda/\partial\rho > 0$ .  $\square$

**Proof of Corollary 2.** The first part of the statement stems from the fact that  $\Delta\lambda^* - \Delta\lambda$  of Eq. (8) is negative since  $\Delta\lambda^* = 2(\rho^*/\rho) + (M-1)\rho^*/2 + (M-1)\rho^*\sqrt{\sigma_p^2 - \rho\sigma_v^2}/\sigma_p\Delta\lambda$ ,  $\Delta\lambda > 0$  (see the proof of Corollary 1),  $\sigma_p^2 > \sigma_v^2$ ,  $\rho^* \leq \rho$ ,

and  $\sqrt{\sigma_p^2 - \rho\sigma_v^2}/\sigma_p < 1$ . Furthermore, it can be shown that  $(\partial\Delta\lambda^* - \Delta\lambda)/\partial\sigma_p > 0$  for  $\partial\rho^*/\partial\sigma_p = 2\rho\sigma_p\sigma_v^2(1-\rho)/(\sigma_p^2 - \rho\sigma_v^2) \geq 0$ ,  $\partial\sqrt{\sigma_p^2 - \rho\sigma_v^2}/\partial\sigma_p = \rho\sigma_v^2/\sigma_p^2 \sqrt{\sigma_p^2 - \rho\sigma_v^2} > 0$ , and  $\partial\Delta\lambda/\partial\sigma_p = 0$ . Lastly,  $\lim_{\sigma_p \rightarrow \infty} \rho^* = \rho$  and  $\lim_{\sigma_p \rightarrow \infty} \Delta\lambda^* = \Delta\lambda$ .  $\square$

**Proof of Remark 2.** The statement stems from the fact that  $\partial\Delta\lambda^* - \Delta\lambda/\partial\rho$  can be shown to be a complex rational function of  $\rho$  whose highest non-negative-integer power in the numerator (denominator) is 4 (2) and whose critical values are complex functions of  $M$ . In particular, algebraic analysis of  $\partial\Delta\lambda^* - \Delta\lambda/\partial\rho$  shows that there exists only one stationary value  $\rho \in (0, 1]$  for  $\Delta\lambda^* - \Delta\lambda$  when  $M$  is either large or small ( $M = 2$  or  $3$ , as in the proof of Remark 1), and an additional critical (either stationary or inflection, depending on  $M$  and  $\sigma_p^2$ ) value otherwise.  $\square$

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