



ELSEVIER

Journal of Banking & Finance 25 (2001) 597–612

**Journal of
BANKING &
FINANCE**

www.elsevier.com/locate/econbase

A note on price noises and their correction process: Evidence from two equal-payoff government bonds

B. Lauterbach ^{*}, A. Wohl

School of Business Administration, Bar-Ilan University, Ramat Gan 52900, Israel

Received 18 January 1999; accepted 20 December 1999

Abstract

The study offers the most direct evidence to date on price noises in call auctions and their correction. We examine a unique sample of two identical securities (two equal-payoff Israeli government bonds) that were traded on separate yet almost simultaneous auctions on the Tel-Aviv Stock Exchange (TASE). The prices of the bonds were equal on average. However, on most of the sample days there were price differences between the bonds. Various estimates suggest that the price noise in one bond is practically uncorrelated with that of the other, and both disappear by the end of the next-day auction. © 2001 Elsevier Science B.V. All rights reserved.

JEL classification: G14

Keywords: Market microstructure; Market efficiency; Price noise; Price correction

^{*} Corresponding author. Present address: The Anderson School at UCLA, 110 Westwood Plaza, Los Angeles 90095, USA. Tel.: +1-310-2066801; fax: 1-310-2065455.

E-mail addresses: beni.lauterbach@anderson.ucla.edu, lauteb@popeye.cc.biu.ac.il (B. Lauterbach), wohlau@mail.biu.ac.il (A. Wohl).

1. Introduction

It is widely accepted that transaction prices include noise. The price noise has many possible sources including temporary liquidity pressures (on the demand or supply side), bid–ask spreads, difficulties and limitations on short sales, discreteness of prices, market segmentation, and the activity of some irrational traders. Hence, price noise is an integral component of security prices.

This study focuses on the price noises in daily call auctions. We examine the prices of two identical-payoff government bonds that traded in almost simultaneous single-daily auctions on the Tel-Aviv Stock Exchange (TASE). Our findings contribute to the literature in several ways. First, we identify a random (mean zero) price difference between the bonds. The mean absolute value of the difference is about 0.2%. This is direct evidence of pricing noise since both bonds should have had identical prices. Previous literature such as French and Roll (1986), Amihud and Mendelson (1987, 1991), Stoll and Whaley (1990), and Hasbrouck (1993) did not observe the pricing noise directly. Rather, it inferred the existence of pricing noise from the fact that subsequent prices appear to exhibit a correction process. It must be noted that the price noises do not imply investor irrationality. We do not find any trading rule that can offer abnormal returns net of transaction costs.

Second, we document a rational correction process. If both bond prices are equally noisy, then a simple average of bond prices is the best estimate of bond value. We present evidence suggesting that the correction process is swift (happens on the next trading day), rational (yesterday's lower price bond tends to increase more today), and relies on the simple average of bond prices. Previous studies did not provide such explicit tests on how rational the correction process is.

Finally, we employ official TASE order imbalance (excess demand/supply) data to investigate the role of order imbalance in creating and correcting the price noise.¹ We show how order imbalances help close the previous day price gap, and find that unexpected order imbalances play some (very moderate) role in creating price noises.

The paper is organized as follows. Section 2 presents a model of noisy prices and discusses its empirical implications. Section 3 describes the trading environment and data. Sections 4 and 5 report the empirical results, and Section 6 concludes.

¹ Order imbalance is the pre-trading (“opening”) excess demand/supply in the security, calculated by the TASE at the previous day closing price – see Section 3.1.

Table 1
Descriptive statistics for the sample bonds in the period June 1994–May 1995

	Mean	S.D.	Minimum	Maximum
<i>Bond 1^a</i>				
Bond price (in NIS, per 100 NIS face value) ^b	159.35	4.22	148.40	167.90
Daily trading volume (NIS) ^b	86,035	145,800	0	1,532,230
Number of trading days	203			
<i>Bond 2^c</i>				
Bond price (in NIS, per 100 NIS face value) ^b	159.29	4.22	147.90	164.60
Daily trading volume (NIS) ^b	68,967	107,416	0	621,200
Number of trading days	177			

^a Name: Sagi 4119; 3% coupon bond; maturity 31 January 1996; fully linked to CPI. Total issued face value: 29,972,075 New Israeli Shekels.

^b NIS is the Israeli currency (New Israeli Shekels). During the sample period \$1 \equiv 3 NIS.

^c Name: Sagi 4120; 3% coupon bond; maturity 31 January 1996; fully linked to CPI. Total issued face value: 20,774,520 New Israeli Shekels.

2. A model of price errors and their correction

2.1. Model preliminaries and assumptions

Consider two equal payoff bonds denoted as 1 and 2 hereafter. Let $P1_t$ and $P2_t$ be the prices of bond 1 and bond 2 on day t , and let V_t be the true value of these bonds. We postulate that the true values of these bonds are identical because these are equal payoff bonds that are traded on the same market and have approximately the same trading volume (see Table 1).²

Denote the natural logarithms of the above variables by $p1_t$, $p2_t$ and v_t , respectively.³ We assume that:

$$p1_t = v_t + e1_t \quad \text{and} \quad p2_t = v_t + e2_t, \quad (1)$$

where $e1_t$ and $e2_t$ are normally distributed, mean zero, and serially uncorrelated random variables with standard deviations of σ_1 and σ_2 , respectively. The terms $e1_t$ and $e2_t$ represent the price noises in the daily auctions of these bonds.

² The similarity of trading methods and volumes is important because liquidity and trading procedures influence security value (see Amihud and Mendelson, 1986; Brennan and Subrahmanyam, 1996; Amihud et al., 1997).

³ The natural logarithm formulation is common in microstructure studies – see, for example, Amihud and Mendelson (1987, 1991), and Stoll and Whaley (1990). This formulation reduces the deviation of the return from the Normal distribution, and is more tractable. Because the absolute bond returns in the sample are relatively small, the empirical results and conclusions are virtually identical when prices (P_t) are used in place of natural logarithms of prices (p_t).

It is further common and convenient to assume that the true value of the bonds (v_t) develops as

$$v_t = v_{t-1} + d + \delta_t, \quad (2)$$

where d is a drift term, and $\delta_t \sim N(0, V_\delta)$. We assume that δ_t is uncorrelated with the price noises, that is, for every t and j $\text{COV}(\delta_t, e1_j) = 0$ and $\text{COV}(\delta_t, e2_j) = 0$.

2.2. The implicit value inference

Given the bond prices, what is the "best" estimate of their value? Under the structure detailed above, a weighted average of $p1_t$ and $p2_t$ of the form $w_1 p1_t + (1 - w_1) p2_t$ is an unbiased estimate of v_t . Further, an unbiased estimate that has minimum square prediction errors can be derived by

$$\text{Min}_{w_1} E [w_1 p1_t + (1 - w_1) p2_t - v_t]^2, \quad (3)$$

where w_1 is the weight of the price of bond 1 in the implicit value equation. The solution of the minimization problem is

$$w_1 = (2\sigma_2^2 - \sigma_{12}) / (2(\sigma_1^2 + \sigma_2^2 - \sigma_{12})), \quad (4)$$

where σ_{12} is the covariance between the pricing errors of the bonds, i.e., $\text{COV}(e1_t, e2_t)$.

Eq. (4) indicates that the optimal weights of bond 1 (w_1) and bond 2 ($1 - w_1$) depend on their noise variability. The bond with the higher price precision (lower noise variability relative to the other bond) is more heavily weighted when inference about v_{t-1} is sought.

How do we estimate w_1 ? We can rewrite the price formation equations as

$$p1_t = d + w_1 p1_{t-1} + (1 - w_1) p2_{t-1} + \varepsilon1_t,$$

$$p2_t = d + w_1 p1_{t-1} + (1 - w_1) p2_{t-1} + \varepsilon2_t.$$

Upon rearranging, we obtain the following "error correction" equations:

$$p1_t - p1_{t-1} = d + (1 - w_1)(p2_{t-1} - p1_{t-1}) + \varepsilon1_t, \quad (5)$$

$$p2_t - p2_{t-1} = d + w_1(p1_{t-1} - p2_{t-1}) + \varepsilon2_t. \quad (6)$$

An efficient way to fit the error correction process is to view it as a multivariate regression system:

$$p1_t - p1_{t-1} = \alpha_1 + \beta_1(p2_{t-1} - p1_{t-1}) + v1_t, \quad (7)$$

$$p2_t - p2_{t-1} = \alpha_2 + \beta_2(p1_{t-1} - p2_{t-1}) + v2_t, \tag{8}$$

and use seemingly unrelated regressions (SUR) for the estimation.⁴

2.3. Testable implications

2.3.1. Equality of prices

The first testable implication is that bond prices are equal on average, that is $E(P1_t - P2_t) = 0$. This implication follows the assumption that the $e1_t$ and $e2_t$ terms in Eq. (1) are mean zero. The bonds have identical values and their price errors are random. Hence, their prices should be equal on average.

2.3.2. Fast correction process

The second testable implication is that $COV(p2_t - p1_t, p2_{t-1} - p1_{t-1}) = 0$. If price noises are mean zero and serially uncorrelated, i.e., vanish quickly (by the next trading day),

$$COV(p2_t - p1_t, p2_{t-1} - p1_{t-1}) = COV(e2_t - e1_t, e2_{t-1} - e1_{t-1}) = 0.$$

In contrast, if the price noises ($e1$ and $e2$) are serially or cross-correlated, the above COV would not equal zero in general.

2.3.3. Special case: Price noises are of equal variance

According to Eqs. (7) and (8), if both bonds have noises of equal variance, then in the multivariate regression system

$$p1_t - p1_{t-1} = \alpha_1 + \beta_1(p2_{t-1} - p1_{t-1}) + \varepsilon1_t,$$

$$p2_t - p2_{t-1} = \alpha_2 + \beta_2(p1_{t-1} - p2_{t-1}) + \varepsilon2_t,$$

both slope coefficients should equal 1/2, i.e., $\beta_1 = \beta_2 = 1/2$. This is an intuitively appealing prediction. For if both bond 1 and bond 2 prices are equally noisy, the best estimate of correct value is their average price.

3. Background and data

3.1. The bonds and trading environment

On January 1991, due to some administrative problems during the Gulf War, the Israeli Treasury issued two identical-payoff bond series: Sagi 4119 and

⁴ Judge et al. (1988, ch. 11) show that SUR provides more efficient estimators than other least squares methods, in the case of cross-equation parameter restrictions.

Sagi 4120.⁵ These bonds are denoted as bond 1 and bond 2, respectively, hereafter. Both bonds matured 31 January 1996, were fully linked to the Consumer Price Index, and paid a yearly coupon of 3% (which was fully indexed as well). The total issued face value of these bonds were 30.0 million New Israeli Shekels (NIS) of bond 1 and 20.8 million NIS of bond 2. (During the sample period \$ 1 \cong 3 NIS.)

The government bond auction had identical principles throughout the sample period. It proceeded as follows. In the morning, public investors and professional traders submitted orders to the TASE computer. Then, TASE published and propagated via computer networks the “order imbalance”. This is the excess demand or supply for each bond, calculated at the previous day closing price. Next, time was given to the public and professional traders to submit offsetting orders. Only orders against the published order imbalance gap could be submitted at this stage. For example, if the order imbalance indicated excess demand (at the previous day’s closing price) for the bond, only sell orders (or cancellation of buy orders) were allowed in this second bidding stage. At the end of the offsetting orders stage, the equilibrium auction price was determined and all possible transactions were executed at this price. More details on the trading mechanism can be found in Bronfeld (1995). Till 17 November 1994 the bonds had back to back auctions on the exchange floor (bond 1 first). On November 20, 1994 the TASE moved to computerized bond trading, and all government bonds had their single daily auction simultaneously.⁶

These bonds were typically held and traded by institutional investors such as pension funds and mutual funds. Short selling of these bonds was not allowed, and there were no “strips” and derivatives.

3.2. *Data*

Daily data for the bonds in the sample were obtained from the TASE. The file includes closing prices, volumes of trade and opening order imbalances. Our sample starts on June 1994. Before June 1994 bonds were traded using a peculiar mechanism: bonds were divided into groups, and the daily auction fixed the same Yield-to-Maturity to all of the bonds in each group. We could

⁵ From our conversations with several key administration officers at the time, we can construct the following (unconfirmed) story. On January 1991 the Treasury started a new auction system: discriminatory price auctions. In fact, Sagi 4119 was the first bond issued using discriminatory price auctions. When the need for further financing emerged on the same month, the Treasury was unsure on whether or not it could legally issue more of the same bond (Sagi 4119), and decided to issue a new bond, Sagi 4120, which was identical to Sagi 4119.

⁶ We sought but could not find significant differences between these subperiods.

not get data on this period. The sample ends in May 1995 because in later months (and up until January 1996, the maturity date) trading in the bonds became extremely thin. For example, on June 1995, bond 2 did not trade at all in half of the month trading days.

Table 1 presents descriptive statistics for the bonds. The average and standard deviation of the price of bond 1 are indistinguishable from those of bond 2. The average daily volume of bond 1 is however 25% higher, probably because of the larger issued quantity of bond 1 (30 million NIS face value compared to 20.8 million NIS face value of bond 2 – a ratio of 1.44). The slightly better tradability of bond 1 is also manifested by the number of active trading days – 203 in bond 1 compared to 177 in bond 2 (see Table 1).

4. Empirical results

4.1. Price differences between the bonds

The two bonds in the sample have identical payoffs. Hence, the first testable implication of the study was that bond 1 and 2 prices would equal, on average (see Section 2.3.1). Table 2 describes the distribution of the price difference between the bonds. The mean price difference is small (3.78 basis points) and insignificantly different from zero (at the 10% level). Furthermore, the median price difference is zero, and price differences appear random: on 102 days the price of bond 1 was higher, on 97 days it was lower, and on 41 days it was equal to that of bond 2. Thus, overall, the equal average price hypothesis is upheld by the data.

It is interesting that on most of the sample days (199 out of 240) the prices of the bonds differ. To gain some perspective on the fluctuations about a zero price difference, we calculate the absolute price difference between the bonds. As reported in Table 2, The average absolute price difference is about 23 basis points (0.23%), and the median absolute price difference is about 13 basis points.

One of the possible reasons for the documented price inequality is that on some of the sample days only one bond was traded. On such a day, the price of the nontraded bond was not updated, and a wedge appeared between the prices of the bonds. To avoid this bias, the bottom part of Table 2 examines only days on which both bonds traded. The results appear similar to those of the overall sample. The price difference between bonds 1 and 2 is random, its mean is insignificantly different from zero, and its median is zero. The proportion of days on which the price of bond 1 was higher (out of the days of unequal price) is 46.3%, which is insignificantly different from 50% (p -value of 0.42) according to the binomial test. The average absolute price difference is 20.5 basis points (median is 12.6 basis points).

Table 2
Characterizing the price differences between the bonds

	Mean	Median	S.D.	Min.	Max.	Num \neq 0	Num $>$ 0
<i>All trading days in the sample (n = 240)</i>							
Price difference between bonds 1 and 2 (in basis points) ^a	3.78	0	37.5	-105	290	199	102
Absolute price difference (in basis points)	-23.3	13.2	29.5	0	290	199	199
<i>Days on which both bonds traded (n = 152)</i>							
Price difference between bonds 1 and 2 (in basis points) ^a	-1.22	0	30.6	-88	99	121	56
Absolute price difference (in basis points)	20.5	12.6	22.7	0	99	121	121
<i>Days on which both bonds traded for the second day consecutively (n = 104)</i>							
Price difference between bonds 1 and 2 (in basis points) ^a	-3.71	0	29.0	-88	99	80	34
Absolute price difference (in basis points)	19.5	12.6	21.9	0	99	80	80

^a Estimated as: $10,000 * (P1 - P2)/(P1 + P2)/2$.

The bottom part of Table 2 presents tests of the price equality hypothesis in the main subsample of the study. This subsample consists of 104 days on which both bonds traded for the second day consecutively. Our price correction relations, derived in Section 2, presume that both bonds were traded on both day t and day $t - 1$. Hence, the 104 days subsample, which satisfies this presumption, is the sample employed in the rest of the study. The behavior of the 104 days sample resembles that of the larger samples. The average price difference between the bonds is insignificantly different from zero, and the mean absolute price difference is 19.5 basis points. Thus, the hypothesis of a mean zero random price difference between the bonds is clearly supported.⁷

⁷ It is interesting that the price difference between bonds 1 and 2 is positive, on average, when all 240 sample days are considered, and negative, on average, on trading days. This is probably due to the fact that bond 1 traded on more days, hence, more often bond 2 price lagged behind (the mean return of the bonds was positive). Anyway, the key finding remains that in all attempted classifications of the data, the mean price difference between the bonds is insignificantly different from zero.

4.2. Swift correction process

Next, we have examined the second elementary testable implication of the model, i.e., that price differences are serially uncorrelated (see Section 2.3.2). In the 104 days sample (where both bonds have trading volume and both have positive volume at the previous day), the first-order autocorrelation of the price differences is 0.007 only (*p*-value of 0.94). This cannot reject the hypothesized swift noise correction process.

4.3. Price correction

It is likely that the price differences between the bonds evolve because of microstructure reasons: each bond has a separate single daily auction with no further trade after it. Price correction occurs on the next trading date. In Section 2 we have shown that the pattern of the price correction depends on the properties of the price noises. For example, if the noises are of equal variability, price correction may use a simple average of yesterday's bond prices as an estimate of yesterday's true value.

To test which price correction process is most consistent with the data, or alternatively which properties of the noise are evident, the multivariate regression system

$$p1_t - p1_{t-1} = a_1 + b_1(p2_{t-1} - p1_{t-1}) + \varepsilon1_t,$$

$$p2_t - p2_{t-1} = a_2 + b_2(p1_{t-1} - p2_{t-1}) + \varepsilon2_t$$

is estimated, using SUR, in three ways:

1. unrestricted;
2. assuming $a_1 = a_2$ and $b_1 = 1 - b_2$, as is implied by the error correction equations (5) and (6);
3. assuming $a_1 = a_2$ and $b_1 = b_2 = 1/2$ (this pattern emerges when noises are of equal variance – see Section 2.3.3).

Table 3 reports the results of the tests. In the unrestricted system the coefficient b_1 is slightly higher than 1/2, and the coefficient b_2 is slightly lower than 1/2. Recalling Eqs. (4)–(8), this finding suggests that the noise in bond 1 price is somewhat higher than that of bond 2. Nevertheless, the difference between b_1 and b_2 is statistically insignificant. Thus, the proposition of equal noise variability cannot be rejected. In addition, the sum of b_1 and b_2 is close to 1, indicating that the restriction in (2) above is probably consistent with the data.

Table 3 confirms that restriction (2) is not rejected by the multivariate tests. This assures that our model and its basic implications such as equal drift are consistent with the data. We have also tested the restriction $a_1 = a_2$ alone, and found that it cannot be rejected at the 5% level. Last, the most extreme

Table 3
The price correction process^a

Restrictions	Parameter estimates (S.E.)				Chi-square likelihood ratio test of the restrictions (degrees of freedom, <i>p</i> -value) ^b
	<i>a</i> ₁	<i>a</i> ₂	<i>b</i> ₁	<i>b</i> ₂	
None	0.00039 (0.00033)	0.00086 (0.00031)	0.53 (0.10)	0.46 (0.10)	–
Identical drift <i>a</i> ₁ = <i>a</i> ₂ and <i>b</i> ₁ = 1 – <i>b</i> ₂	0.00065 (0.00028)	0.00065 (0.00028)	0.54 (0.09)	0.46 (0.09)	2.73 (2, 0.26)
Identical drift and identical price noise variances <i>a</i> ₁ = <i>a</i> ₂ and <i>b</i> ₁ = <i>b</i> ₂ = 1/2	0.00065 (0.00028)	0.00065 (0.00028)	0.50	0.50	2.91 (3, 0.41)

^aThe equation system

$$r_{1,t} = p1_t - p1_{t-1} = a_1 + b_1(p2_{t-1} - p1_{t-1}) + \varepsilon1_t,$$

$$r_{2,t} = p2_t - p2_{t-1} = a_2 + b_2(p1_{t-1} - p2_{t-1}) + \varepsilon2_t,$$

where *p*₁ and *p*₂ are the natural logarithms of the prices of bonds 4119 and 4120 (respectively), is estimated simultaneously using the SUR methodology. Then, several restrictions on the system are tested. The sample consists of all days on which both bonds traded for the second day consecutively (*n* = 104).

^bThe likelihood ratio test statistic is developed and discussed in Gallant and Jorgenson (1979). In large sample it is distributed chi-square with *R* degrees of freedom, where *R* is the number of restrictions imposed on the system.

restriction (3) is also found consistent with the data.⁸ The chi-square likelihood ratio test statistic of restriction (3) is 2.91, and its *p*-value is 0.41. It appears that the bond price noises may have equal variances.

The findings in Table 3 can also be summarized from the price correction perspective. The market seems to understand the price noises generated by auctions. When it (the market) encounters two prices, which are essentially two equally noisy estimates of correct value, it uses an average of these two prices to rationally assess the true value.

5. Extensions and further results

5.1. Estimating true value and noise variabilities

Previous research expressed interest and estimated the magnitude of the pricing noise (its standard deviation). Amihud and Mendelson (1987), and Stoll

⁸ This is not surprising given the fact that the unrestricted estimates of *b*₁ and *b*₂ (0.53 and 0.46, respectively in Table 3) are close to 1/2.

and Whaley (1990) compare the open-to-open and close-to-close stock return variabilities on the New York Stock Exchange. They find that the open-to-open return variances of stocks are on average 20% larger than the close-to-close return variances. This difference suggests that opening prices are noisier than closing prices. Closing prices include noise as well. French and Roll (1986) use variance ratios to assess that on average 4–12% of the close-to-close return variance (of NYSE and AMEX stocks) is caused by mispricing. Hasbrouck (1993) examines intradaily transaction prices of a sample of NYSE stocks, and estimates that the average standard deviation of the intradaily price noise is 0.33% (of stock price).

While previous research focused on equity trading noise, this study offers a view on bond price noise. We can use the basic model of Section 2 to estimate noise variability. Let us define the bond returns on day t :

$$r1_t = p1_t - p1_{t-1}, \quad r2_t = p2_t - p2_{t-1}.$$

From Eqs. (1) and (2),

$$\text{VAR}(r1_t) = \text{VAR}(p1_t - p1_{t-1}) = \text{VAR}(\delta_t) + 2\text{VAR}(e1_t), \tag{9}$$

$$\text{VAR}(r2_t) = \text{VAR}(p2_t - p2_{t-1}) = \text{VAR}(\delta_t) + 2\text{VAR}(e2_t), \tag{10}$$

$$\text{COV}(r1_t, r2_t) = \text{VAR}(\delta_t) + 2\text{COV}(e1_t, e2_t), \tag{11}$$

$$\text{COV}(r1_t, r2_{t-1}) = \text{COV}(r2_t, r1_{t-1}) = -\text{COV}(e1_t, e2_t). \tag{12}$$

The point estimates of the left-hand side variances and covariances in Eqs. (9)–(12) are:

$$\text{VAR}(r1_t) = 1379 \times 10^{-8},$$

$$\text{VAR}(r2_t) = 1204 \times 10^{-8},$$

$$\text{COV}(r1_t, r2_t) = 367 \times 10^{-8},$$

$$\text{COV}(r1_t, r2_{t-1}) = 93 \times 10^{-8},$$

$$\text{COV}(r2_t, r1_{t-1}) = 24 \times 10^{-8}.$$

The last two numbers above are both estimates of $-\text{COV}(e1_t, e2_t)$. Taking their average, the point estimate of $\text{COV}(e1_t, e2_t)$ is -58.5×10^{-8} . Plugging this estimate into (11) yields a $\text{VAR}(\delta_t)$ estimate of 584×10^{-8} . Solving (9) and (10) we get $\text{VAR}(e1_t) = 398 \times 10^{-8}$ and $\text{VAR}(e2_t) = 310 \times 10^{-8}$. It appears that both price noises as well as the fundamental value innovation (δ_t) have a standard deviation in the order of magnitude of 0.2%.

The price noise standard deviation of about 0.2% is consistent with the average absolute price difference of about 0.2%. The estimated noise standard

deviation also appears reasonable given: (1) the minimum tick size (0.06–0.07% based on the average price of the bonds); (2) the one-way commission of about 0.15% charged from institutional traders and professional portfolio managers; and (3) the fundamental value (δ_t) standard deviation of 0.2%. (We speculate that if $\text{VAR}(\delta_t)$ was considerably higher, noise variability would increase as well.)

5.2. Testing for excess return opportunities

The evidence so far suggests a rational bond trading market with intelligible inference of true value and swift noise correction. This rational image can be further challenged by arbitrage-type tests. The search for arbitrage schemes is limited, however, by one institutional factor. During the sample period, the Bank of Israel banned short sales of government bonds, arguing that short sales interfere with its monetary policy. Therefore, our investigation is limited to strategies that require initial holdings in both bonds.

Three nonshort-sales-based trading rules are attempted, assuming a trader who owns both bonds. The first is based on yesterday's (day $t - 1$) closing prices: if day $t - 1$ prices differ, on day t buy the day $t - 1$ cheaper bond and sell the day $t - 1$ more expensive bond. This trading rule could be implemented in 114 days, and its average return, calculated as the price of the bond sold minus the price of the bond bought divided by yesterday's average price of the bonds, is 0.04% (p -value of 0.16). Given the transaction costs of 0.3% (twice the regular commissions – once for the buy and once for the sell orders), it becomes apparent that this trading rule was unable to generate any economically or statistically significant excess returns.

The second conceivable trading rule is based on order imbalances. If, today, one bond has excess demand and the other excess supply, buy the bond with excess supply (whose price is likely to decline) and sell the bond with excess demand. Recall that order imbalances are published before prices are settled, so that offsetting orders of the type proposed above are accepted by the exchange. This strategy was feasible on 45 days. Its (before transaction costs) average return, 0.09% (p -value of 0.02), is, however, economically unattractive.

Last, we attempted a strategy based on both prices and order imbalances. If bond X has a lower (or equal) price yesterday and an excess supply today, while bond Y has a higher (or equal) price yesterday and an excess demand today, buy bond X and sell bond Y . This strategy could be implemented on 8 days, and its average excess return, 0.38% (p -value of 0.007), is statistically significant. Nevertheless, given transaction costs (0.3%) and possible market impact costs, it becomes unlikely that such a trading rule could generate economically significant excess returns.

In sum, the trading rule tests support the contention that the pricing noises are limited in magnitude. Pricing noises appear bound by the level of transaction costs, and by a variety of feasible trading rules.

5.3. *The role of order imbalance in correcting and creating price distortions*

One of the unique features of our data set is that it includes official TASE order imbalance data. TASE calculated order imbalance based on the orders submitted to it before the auction starts. The order imbalance on day t is measured by TASE as the difference between the aggregate bids to buy the stock and aggregate bids to sell, at the previous day closing prices. In other words, this is the initially revealed excess demand at the previous day's closing price. Because investors are allowed to add or cancel orders (under certain conditions – see Section 3.1) in the second round of bidding before auction price is fixed, the official order imbalance variable is only a proxy of the “true” excess demand at this price.

The excess demand schedule is (weakly) decreasing in price. Thus, we expect price differences on day $t - 1$ to impact order imbalances on day t . More specifically, the prediction is that on day t , order imbalance (demand minus supply) would be higher for the day $t - 1$ cheaper bond.⁹ Also, when price differences are larger, the order imbalance differences between the bonds increase because they are proxies for more distant points on the excess demand schedule.

To examine this prediction we run the regression

$$\text{ordimb1}_t - \text{ordimb2}_t = a + b(p_{2,t-1} - p_{1,t-1}) 10,000 + \varepsilon \text{ordimb}_t, \quad (13)$$

where $\text{ordimb}N_t$ is the order imbalance of bond N on day t . The analysis above suggests a positive coefficient b , which is confirmed by the tests. The estimated coefficient b is 7120 with a t -statistic of 9.9, and the adjusted R^2 is 0.64. Evidently, if on day $t - 1$ one bond is cheaper, it would have a larger order imbalance (relative to the other bond) on the day t auction. This result illustrates the mechanics of price correction. A price difference on day $t - 1$ generates order imbalances on day t that favor the cheaper bond. These order imbalances “power” the price correction process.

It is interesting to examine whether order imbalances have also a role in creating the price differences. One could hypothesize that a relatively large unexpected order imbalance in one of the bonds would distort the price of that bond relative to the other bond. This proposition is based on some inelastic

⁹ Suppose that both bonds have identical demand and supply schedules, and recall that order imbalance is measured at the previous day price. Given these, it is natural that at the lower price of yesterday's cheaper bond, the excess demand is higher.

demand and supply curves for the bonds and/or on some market friction and imperfections of the trading system. It basically suggests that unexpected supply/demand causes an unjustified price movement, i.e., a (temporary) departure of price from value. In short, unexpected liquidity shocks induce noises in prices.

We tested this proposition in a special subsample: following days of equal bond prices. The advantage of this special subsample is that on trading days that follow days of equal price, order imbalances do not depend on the previous day price gap (because previous day price gap is zero). Hence, on such days, the unexpected order imbalance difference between the bonds (which presumably generates the price differences) may be approximated by the simple order imbalance difference between the bonds.

The special subsample includes 23 observations, and the fitted regression is

$$(p1_t - p2_t) 10,000 = -1.7 + 0.000136(\text{ordimb}1_t - \text{ordimb}2_t) + \eta_t.$$

The regression coefficient of the order imbalance difference is marginally statistically significant (its *t*-statistic is 2.0). However, the adjusted R^2 is 0.02 only. Clearly, the exact determinants of the price noise (price differences between the bonds) have not been uncovered.

6. Summary and concluding remarks

The study offers the most direct evidence to date that auction prices include noise. Two equal payoff Israeli government bonds, traded almost simultaneously in separate daily call auctions on the TASE, had equal prices on average. However, on most of the sample days there were price differences between the bonds. These price differences are clearly noise.

Although we find price noises, we do not find evidence of “market irrationality”. Various estimates indicate that the price noise in one bond is practically uncorrelated with that of the other, and both disappear by the end of the next-day auction. A direct search for excess return opportunities (after transaction costs) fails to disclose any.

Finally, order imbalances are shown to propel the price correction process. If the auction price of one bond is lower than that of the other, then the next-day order imbalance acts to close the gap. (There appears a relatively higher demand for the cheaper bond.) We also present evidence that unexpected order imbalances contribute to the price noise. This is consistent with the hypothesis that price noises are caused by supply/demand shocks. However, this relation is relatively weak, and the exact sources of the price noise have not been unveiled.

The study concludes with a policy suggestion. It seems that enabling more flexible limit orders such as “Buy 1500 of the cheapest between bonds 1 and 2” could diminish the noise and improve price efficiency. Introducing complex

orders is also suggested by Amihud and Mendelson (1985, 1988, 1990), Beja and Hakansson (1979), Brown and Holden (1993), Economides and Schwartz (1995), Miller (1991), and Wohl and Kandel (1997). Our case of equal-payoff bonds is a clear example where such orders could be beneficial. More generally, complex orders appear desirable whenever there is a correlation between the traded assets.

Acknowledgements

The comments of Yakov Amihud, two anonymous referees and the Editor (Giorgio Szego) are gratefully acknowledged.

References

- Amihud, Y., Mendelson, H., 1985. An integrated computerized trading systems. In: Amihud, Y., Ho, T., Schwartz, R. (Eds.), *Market Making and the Changing Structure of the Securities Industries*. Lexington Books, Lexington, MA.
- Amihud, Y., Mendelson, H., 1986. Asset pricing and the bid–ask spread. *Journal of Financial Economics* 17, 223–249.
- Amihud, Y., Mendelson, H., 1987. Trading mechanism and stock returns: An empirical investigation. *Journal of Finance* 42, 533–553.
- Amihud, Y., Mendelson, H., 1988. Liquidity volatility and exchange automation. *Journal of Accounting Auditing and Finance* 3, 369–385.
- Amihud, Y., Mendelson, H., 1990. How (not) to integrate the European capital markets. In: Giovannini, A., Mayer, C. (Eds.), *European Finance Integration*. Cambridge University Press, Cambridge.
- Amihud, Y., Mendelson, H., 1991. Volatility efficiency and trading: Evidence from the Japanese Stock Market. *Journal of Finance* 46, 1765–1789.
- Amihud, Y., Mendelson, H., Lauterbach, B., 1997. Market microstructure and securities values: Evidence from Tel Aviv Stock Exchange. *Journal of Financial Economics* 45, 365–390.
- Beja, A., Hakansson, N.H., 1979. From orders to trades: Some alternative market mechanisms. In: Bloch, E., Schwartz, R. (Eds.), *Impending Changes for Securities Markets*. JAI, Greenwich, CT.
- Brennan, M., Subrahmanyam, A., 1996. Market microstructure and asset pricing: On the compensation for illiquidity in stock returns. *Journal of Financial Economics* 41, 441–464.
- Bronfeld, S., 1995. *Trading Systems on the Tel Aviv Stock Exchange*. Tel Aviv Stock Exchange Ltd., Tel Aviv, Israel.
- Brown, D., Holden, 1993. The design of limit orders. Working paper, Indiana University, Bloomington, IN.
- Economides, N., Schwartz, R., 1995. Electronic call market trading. *Journal of Portfolio Management* 21 (3), 10–18.
- French, K., Roll, R., 1986. Stock return variances: The arrival of information and the reaction of traders. *Journal of Financial Economics* 17, 5–26.
- Gallant, A.R., Jorgenson, D.W., 1979. Statistical inference for a system of simultaneous nonlinear implicit equations in the context of instrumental variable estimation. *Journal of Econometrics* 11, 275–302.

- Hasbrouck, J., 1993. Assessing the quality of a security market: A new approach to transaction-cost measurement. *Review of Financial Studies* 6, 191–212.
- Judge, G.G., Hill, R.C., Griffith, W.E., Lutkepohl, H., Lee, T.C., 1988. *Introduction to the Theory and Practice of Econometrics*. Wiley, New York.
- Miller, M., 1991. *Financial Innovations and Market Volatility*, Blackwell, Cambridge, MA (Chapter 11).
- Stoll, H., Whaley, R., 1990. Stock market structure and volatility. *Review of Financial Studies* 3, 37–71.
- Wohl, A., Kandel, S., 1997. Implications of an index-contingent trading mechanism. *Journal of Business* 70, 471–488.