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Improved information shock and price dispersion: A natural experiment in the housing market☆

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ABSTRACT

This study employs data from a natural experiment to assess the effect of an improved price information shock on subsequent real estate transaction price dispersion. While transaction data in the Israeli real estate market had not previously been available to the public, in 2010 an Israeli court ordered the Israel Tax Authority to post all real estate transaction data on its website. We employ all housing transactions in the period prior and subsequent to this event to assess its effect on housing price dispersion. Results provide evidence of significant decrease in the dispersion of quality-adjusted prices. Further, we find evidence that the information shock effect on price dispersion varies with household and market characteristics.

1. Introduction

Economists have long recognized the central role of information in the operation of markets. For example, when information is costly or imperfect, sub-optimal welfare is likely to be attained, and market equilibrium may exhibit price dispersion even for homogeneous goods (see, e.g., Stiglitz, 1985, 1961, respectively).[1] However, while a great deal of theoretical and empirical research has been devoted to understanding the effect of information on prices, only limited empirical work has been undertaken to date on the specific effect of information shocks on price dispersion of goods.

An exception to this trend is a study by Jensen (2007) on the effect of improved information technology shock on price dispersion and welfare in the fishing industry in Kerala, India. According to Jensen (2007), information shock that was associated with the introduction of mobile phone service to fishermen and wholesalers has led to a reduction in price dispersion in the South Indian fisheries sector. (Similarly, Aker, 2010 finds that the introduction of mobile phone services reduced price dispersion across grain markets in Niger.)[2]

A recent experience in the Israeli real estate market serves as a natural experiment for further exploring the effect of information availability on price dispersion in a market of non-homogeneous goods where transactions carry considerable long-term, economic consequences. Specifically, in 2010, an Israeli court ordered the Israel Tax Authority to open its records to the public on all past and current real estate transactions. For the first time, price and other related real estate transaction infor-

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[1] Studies on the role of information in markets are too numerous to cite. See Stiglitz (1985) for a thorough review of the role of information in economic analysis.

[2] Note that, unlike Fama et al. (1969) and many others that followed, we do not focus on the price reaction to new (favorable or unfavorable) information per se; rather, along the lines of Jensen (2007), we focus on the effect of improved information on price dispersion. Prevailing rational explanations for the price dispersion of a given good include the cost of information collection (e.g., Stigler, 1961, Rothschild, 1973, and, more recently, Janssen and Moraga-González, 2004 and Janssen et al., 2005) and consumer heterogeneity in a "clearinghouse" setting (e.g., Salop and Stiglitz, 1977; Varian, 1980 and, more recently, Baye and Morgan, 2001 and Baye et al., 2004b). Empirical studies of price dispersion in particular account for explanations such as the absolute value of the good (e.g., Pratt et al., 1979 and, more recently, Gatti and Kattuman, 2003 and Eckard, 2004); purchase frequency (e.g., Sorensen, 2000); number of competing sellers in the market (Borenstein and Rose, 1994; Baye et al., 2004a; Barron et al., 2004; Lewis, 2008 and Chandra and Tappata, 2011); and search cost (e.g., Marvel, 1976; Brown and Goolsbee, 2002; Brynjolfsson and Smith, 2000; Smith and Brynjolfsson, 2001; Dinlersoz and Li, 2006 and Chandra and Tappata, 2011). Several studies also examine the persistence of price dispersion over time (see, among others, Lach, 2002). Finally, see the comprehensive review of price dispersion literature in Baye et al. (2006a).

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mation was disclosed to market participants and was freely accessed through the Tax Authority's website.[3]

We study the effect of this exogenous information shock on the price dispersion of subsequent housing transactions. In particular, by observing all market transaction prices prior and subsequent to the improved information shock, we estimate the controlled change in the dispersion of quality-adjusted housing prices over time and across locations. We further examine the sensitivity of the estimated price dispersion effect to household socioeconomic (SES) status and housing asset heterogeneity in the market. As our framework only allows for a before/after-type test (as the new information disclosure was implemented nationwide), we conduct a series of identification and robustness checks for our findings. These tests include using rental housing transactions as a control group; changing the scope of the geographical areas for which price dispersion is estimated; quantitatively assessing the degree of potential unobservable variable bias along the lines of Altonji et al. (2005); and applying placebo as well as nonparametric tests.[4]

Our results provide evidence of decreased price dispersion that follows the exogenous public disclosure of transaction price information. Specifically, over various test specifications, we find that standard deviation of quality-adjusted (hedonic) prices decreased by about 18% subsequent to the improved information shock. Moreover, applying the Altonji et al. (2005) method for testing whether unobservable variables can explain the decreased price dispersion effect, we find that the normalized shift in the distribution of the unobservables should be about 2.8 (6.3) greater than the normalized shift in the distribution of the treatment (the observed information shock) to nullify the entire 3- (1-) year pre/post effect of the shock. Placebo and nonparametric tests further support our outcomes on the significant decreased price dispersion that follows the information shock.

In addition, we find that the magnitude of the effect of improved information shock on decreased price dispersion is negatively correlated with household SES level in the market. This is consistent with the notion that market participants in less privileged areas, having limited access to means that may ameliorate information shortage, are experiencing a greater effect of the improved public information. We also find that the magnitude of the improved information shock effect is positively correlated with asset heterogeneity in high SES markets (consistent with Kurlat and Stroebel, 2015—see further discussion in Section 7 below). Our results are robust to a series of sampling and test-design issues.

The intuition underlying the empirical results is that, in a market with imperfect price information (as compared to a market with perfect price information), the seller and the potential buyer may arrive at the negotiation when they observe different subsets of current housing transactions (i.e., they are equipped with different information sets regarding ongoing market prices). Not only does the latter decrease the likelihood of successfully reaching an agreed-upon closing price, but also, since the underlying (imperfect) information sets serving the parties in one transaction may differ from the information sets observed by the parties in another concurrent transaction, different closing prices might attain, *ceteris paribus*—leading to a greater price dispersion under imperfect price information. In the context of our empirical framework, the increased information access provided to market participants by the Israel Tax Authority should thus likely lead to a narrowing of price dispersion among housing transactions.[5]

A number of studies specifically examine the importance of information in the real estate market. Garmaise and Moskowitz (2004) show how market participants tend to purchase nearby assets, trade properties with long income histories, and avoid deals with informed brokers in order to resolve information asymmetries. Levitt and Syverson (2008) further quantify housing price information: informed, as compared to uninformed, sellers (real estate agents versus home sellers, respectively) sell their own properties more patiently, attaining higher prices. Kurlat and Stroebel (2015) further show that the share of informed sellers may predict future price declines, while informed buyers obtain greater price appreciation in the housing market.[6]

The primary contributions of this research are as follows. First, we extend Jensen's (2007) and Aker's (2010) evidence on the price dispersion effect of information shock by not only examining a different type of information shock, but also showing that the effect carries over into a market in which transactions involve considerable personal economic consequences. Moreover, we show that the extent of the decreased price dispersion that follows the improved information shock associates with household and asset characteristics in the market.

The remainder of the paper proceeds as follows: Section 2 provides further institutional background on the Israeli real estate market and the exogenous information shock. Section 3 describes the data and Section 4 describes the basic methodology. Section 5 presents the results and Section 6 presents a series of identification and robustness tests. Section 7 presents the sensitivity of the results to household and market characteristics. Finally, Section 8 provides a summary and concluding remarks.

2. Background

Israeli law requires that the lawyer(s) representing the parties involved in a real estate transaction provide(s) a report to the Israel Tax Authority upon the closing of the transaction.[7] This report must include the closing price as well as information on fundamental attributes of the transacted asset. While the Tax Authority traditionally collected the data on each individual transaction, the information was never available to the public (or to market professionals). Moreover, unlike the multiple listing method employed in the United States, the common practice in Israel is that each broker privately manages his or her listing of assets for sale, which is not shared with other brokers.[8] Thus, in order

[3] Our study thus further relates to the ambiguous evidence on the effect of online markets on price dispersion of goods (see, among others, Bailey, 1998; Brynjolfsson and Smith, 2000; Clemons et al., 2002 and Baye et al., 2006b).

[4] Eerola and Lyytikäinen (2015) use a sample of transactions in the Helsinki (Finland) housing market to explore the effect of improved information shock on the price level and time on the market of transacted assets. Also, in a recent paper, Kotsenko (2018) studies the effect of price information in the Israeli market on price distribution and time on market. Kotsenko, however, uses a different approach—transaction- (rather than market-) level information that is non-shock related—than the one undertaken in our study. Also, Leung et al. (2006) show that price dispersion in the housing market correlates to some extent with macro-economic variables.

[5] Harding et al. (2003) show that hedonic prices may be affected by varying the bargaining powers of the parties to the transaction. In our empirical model, we control for bargaining power that is correlated with up- and down-trends in the price level. Unlike Harding et al. (2003), however, we do not observe characteristics of the sellers and buyers that may correlate with bargaining power. Also, in a one-sided search model, Haurin (1988) shows that "atypicality" of a housing unit (i.e., unique housing characteristics) may affect the variance of its value. While "atypical" characteristics are unobservable in our data, it is reasonable to assume that they distribute approximately uniformly over time, given our extensive transaction dataset. Also, Liu et al. (2016) show that relative housing prices between high- and low-quality segments vary during bust periods. We empirically control for this potential effect by observing the time-varying average price change in every sub-market. Finally, see Han and Strange (2015) for a recent review of the related literature on the effect of real estate market microstructure on prices.

[6] For more on the importance of information in the real estate market, see Downs and Guner (1999). Also, see Chinco and Mayer (2016) on the relationship between information-holding and market timing.

[7] That the lawyers handling the transaction are responsible for the reporting to the Tax Authority considerably decreases any concerns about misreporting.

[8] Housing units for sale in Israel are offered mainly through real estate brokers or person-to-person. (Auctions are very rare.) According to the asking price data available to us, the share of broker transactions in the market is about 58%. Generally, there are a number of brokers dominating each sub-market

to assess market prices, land appraisers, real estate brokers, and other professionals had to rely on limited information sources (such as past transactions in which they were personally involved, and asking prices that had appeared in their listing). Clearly, under such circumstances, the general public (including buyers and sellers) had no formal access to transaction price information—and, in fact, had very limited access to non-formal information sources.

The timeline of events that led to the publication of real estate transaction information in Israel is as follows. In mid-April 2010, following a court order, the Israel Tax Authority begin to publish micro-level information on all real estate transactions in Israel.[9] For the first time, access to nationwide transaction information was provided at no charge through the Tax Authority's website, and the data were continuously updated, with new transaction closings coming in (the original information release included transactions from 1998 onward). In fact, the publication of the data was completed in two phases. In April 2010, the Tax Authority's data website was launched; however, the interface was "unfriendly" and provided only data on the asset price and its location. (Its physical attributes were excluded.) In the second phase, however, some six months later (October 2010), the website was upgraded, allowing for easier access by non-professional users and, in addition, providing more complete information on each transacted asset, including number of rooms, area in square meters, age of the structure, floor number, and the number of floors in the structure where the asset was located.[10]

Fig. 1A–C show statistics of Google Trends' "Search Interest" resulting from a search for the terms "apartment prices tax authority," "real estate information," and "apartment prices" (translated from Hebrew), respectively, over the period January 2007 to December 2014. Several patterns are evident. First, one will note the relatively insignificant search volume up to March 2010. Thereafter, following the media reports that accompanied the launch of the Tax Authority's transaction information website in April, a first search peak is recorded in May 2010. Another peak follows in October-November 2010, concurrent with the upgrade of the Tax Authority website. Finally, in the period subsequent to November 2010, search volume maintains a level greater than that recorded during the pre-information disclosure period, as availability of the data on the web transformed the search for transaction closings into a routine practice for real estate market participants.[11]

Following the above, we hypothesize that the dispersion of closing prices of quality-adjusted assets in the housing market has decreased as a result of the improved information shock. Intuitively, as compared to the situation that prevails under perfect price information, under imperfect price information, the partial information available to the parties in one transaction is likely to differ from the information available to the parties in another concurrent transaction—leading to different closing prices in the two transactions, *ceteris paribus*; hence a greater price dispersion. In the context of our empirical framework, the public disclosure of price information by the Israel Tax Authority should lead to a decreased price dispersion among housing transactions.

3. Sample description

We use the above-described institutional development in the Israeli real estate market to study the effect of the price information shock on housing unit price dispersion. Our sample includes the universe of all secondary market (non-"first-hand") condominium apartment transactions in Israel over the period 2007–2013, a total of 222,163 observations.[12] Specifically, as further described in the next section, we estimate and compare the dispersion of quality-adjusted prices over the three years prior to the price information disclosure of April 2010, when the Tax Authority's website was originally launched with partial information on housing transactions and the three years subsequent to the complete information provision—that is, when the website was upgraded in October 2010. (In that regard, our framework may be considered as a pre/post natural experiment.)

The sample comprises the information that is provided on the Israel Tax Authority website on each housing unit transaction, including the closing price and date as well as a series of asset attributes. Table 1 presents summary statistics of the sample of transactions (all sample and sample stratified by pre- and post-information shock). As indicated in the table, the typical dwelling unit is a 3- to 4-room, 856-square-foot condominium apartment located in a 30-year-old structure. The average unit price is about $248,000, with a standard deviation of about $171,000.[13] It further follows from Table 1 that the average and standard deviation of asset attributes are similar over the period before and after the information shock (see analysis below for the controlled before/after difference in the standard deviation of the price).

Table 2 further presents summary statistics for the sample of city panel (all sample and sample stratified by pre- and post-information shock). As shown, the average value of *Treatment*, an indicator variable that equals 1 for post-information shock periods and zero otherwise, is about 0.5. The table also provides information on a set of control variables, including the 6-month rate of change in quality-adjusted housing prices in city c (denoted by $\Delta\hat{P}_{tc}$), the average of which is 0.05;[14] the

(neighborhood); however, information is rarely shared among brokers. Also, it should be noted that about 95% of the Israeli residential market consists of condominium apartments, where the ownership rate is about 68%, and about 93% of the households are located in urban areas (see Central Bureau of Statistics, 2015).

[9] The disclosure of the data may be considered an *exogenous* shock to the real estate market since, until the judge's order in February 2010, the petition received insignificant media coverage and public attention. The judge's decision was greeted with great surprise and led to the first publication of the data in April 2010—an event that received extensive media coverage. (For proper disclosure: One of the authors of this article was among those who filed the petition to court.)

[10] It should be further noted that in March 2011, another (private) website was launched that provided access to the Tax Authority transaction data at no charge and in a highly accessible and user-friendly manner. Several other private websites followed over the years.

[11] As mentioned above, a number of private websites were launched in the year following the data disclosure by the Tax Authority. This may explain not only the overall search volume post-October 2010, but also the relatively higher volume for the terms "real estate information" and "apartment prices," as compared to "apartment price *tax authority*" in the post-2012 period.

[12] As condominium apartments are about 95% of housing assets in Israel (see Israel Central Bureau of Statistics, 2015), we specifically focus on that segment of the market. We omit primary market ("first-hand") transactions, as they are initiated by developers who, arguably, maintain an information advantage over the reasonable player in the market. Also, in order to reliably assess the price dispersion, we omit observations in cases where fewer than 20 transactions occur in a given city in a given period (quarter); moreover, we require that each city appear at least in one period both prior and subsequent to the information shock (see further description in Section 4 below). Hence, from the raw sample of 284,214 observations over the period 2007–2013, we are left with a final sample of 222,163 observations in 57 cities (out of the 76 cities in Israel), of which 45 appear in all 24 examined quarters and 53 in at least 18 periods. Our estimation outcomes, however, are robust to changing the condition for city/quarter participation in the estimation from 20 transactions to either 30, 40, or 50 transactions (these results are not reported but are available on request).

[13] For convenience, prices are presented in US dollars, where 1 US dollar equals 3.77 New Israeli shekels (NIS).

[14] Note the relatively high average return on the quality-adjusted prices over the examined period. Following Pratt et al. (1979), Gatti and Kattuman (2003), Eckard (2004), and others who present evidence of an inverse correlation between the value of a good and its price dispersion, we control for the changes in quality-adjusted housing price. However, as the unit-root hypothesis for the quality-adjusted price *level* is not rejected (Fisher-type test based on ADF p-value equal to 0.909), we specify this non-stationary control variable in difference terms (the latter is found to be stationary; unit-root hypothesis is rejected at the 1% level). As shown below, our results on the correlation between price returns and price dispersion are consistent with previous literature.

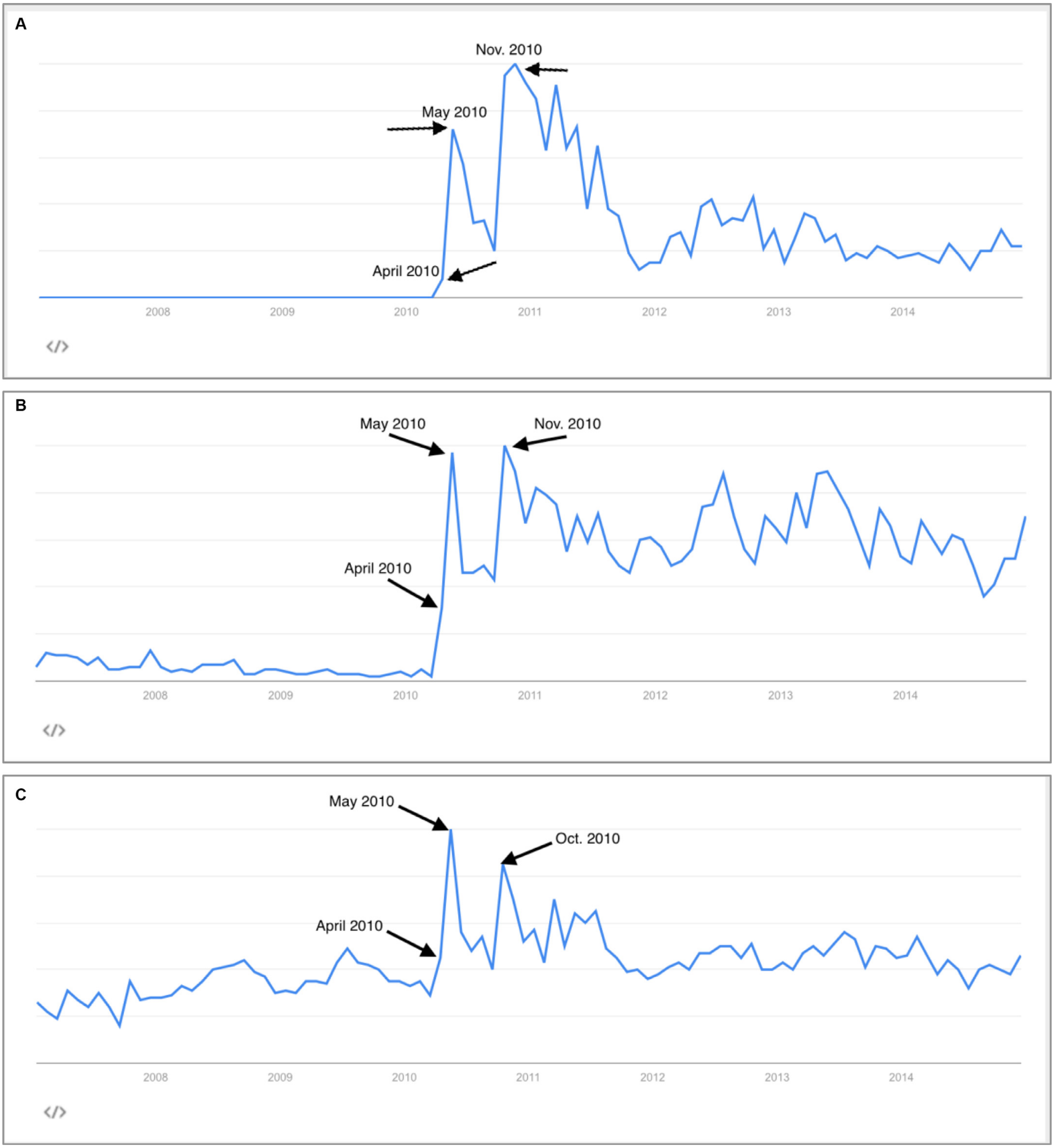

Fig. 1. (A). Google Trend's Search Interest in "Apartment Prices Tax Authority," January 2007 to December 2014. (B). Google Trend's Search Interest in "Real Estate Information," January 2007 to December 2014. (C). Google Trend's Search Interest in "Apartment Prices," January 2007 to December 2014
Notes: The y-axis represents search interest popularity. The peak attains a value of 100, and all levels are relative to the highest point on the chart. A score of 0 means the term was less than 1% as popular as the peak.

number of transactions per quarter t in city c (N_{tc}), the average of which is 173; and quarterly standard deviation of daily yields on the Tel Aviv 100 stock index (the Israeli equivalent of the S&P 500) (SD_Stock_t), the average of which is about 1.23%. In addition, as shown in Table 2, the panel analysis controls for the average of dwelling unit characteristics in city c at time t (including Avg_Area_{tc}, Avg_Rooms_{tc}, Avg_Age_{tc}, and Avg_SES_{tc}) and the variance of dwelling unit characteristics in city c at time t (including SD_Area_{tc}, SD_Rooms_{tc}, SD_Age_{tc}, and SD_SES_{tc}).[15]

[15] Variables for which we could not reject the non-stationarity assumption are included in first difference terms (see variables denoted with Δ). Also, informa-

Table 1
List of variables, description, and summary statistics of sale transactions.

Variable	Sample	All		Pre-shock		Post-shock	
	Description	Avg.	Std.	Avg.	Std.	Avg.	Std.
P	Transaction closing price (in USD)	248,338	171,557	209,862	140,632	286,645	190,008
Area	Floor area (in square feet)	856.2	309.4	855.8	309.7	856.8	309.1
Age	The age of the structure (in years) at the time of the transaction	30.3	16.2	29.1	15.8	31.6	16.6
Rooms_2	Dummy variable that equals 1 if total number of rooms is equal to 1.5 or 2; zero otherwise	0.139	0.346	0.14	0.35	0.13	0.34
Rooms_3	Dummy variable that equals 1 if total number of rooms is equal to 2.5 or 3; zero otherwise	0.430	0.495	0.43	0.50	0.43	0.49
Rooms_4	Dummy variable that equals 1 if total number of rooms is equal to 3.5 or 4; zero otherwise	0.306	0.461	0.30	0.46	0.31	0.46
Rooms_5	Dummy variable that equals 1 if total number of rooms is equal to 4.5 or 5; zero otherwise	0.105	0.306	0.10	0.30	0.11	0.31
Rooms_6	Dummy variable that equals 1 if total number of rooms is equal to 5.5 or 6; zero otherwise	0.019	0.136	0.02	0.14	0.02	0.13

Notes: Table 1 presents summary statistics of the sample of sale transactions (all sample and sample stratified by pre- and post-information shock).

Table 2
List of city-level panel variables, description, and summary statistics.

Variable	Sample	All		Pre-shock		Post-shock	
	Description	Avg.	Std.	Avg.	Std.	Avg.	Std.
$Treatment_t$	Dummy variable that equals 1 for periods subsequent to information disclosure (i.e., subsequent to 2010Q3)	0.52	0.50	0	0	1	0
$SD\_\hat{P}_{tc}$	The 6-month (ending at *t*) standard deviation of quality-adjusted housing prices in city *c*	0.04	0.03	0.05	0.03	0.04	0.02
$\Delta\hat{P}_{tc}$	2-quarters (ending at *t*) rate of change in quality-adjusted housing prices in city *c*	0.05	0.06	0.06	0.07	0.04	0.05
N_{tc}	The number of transactions in period *t* and city *c*	173	184	178	189	168	179
SD_Stock_t	The 3-month moving standard deviation of daily yields of the Tel-Aviv100 stock index	1.23	0.61	1.52	0.65	0.96	0.42
Avg_Area_{tc}	The average area (in square feet) of assets transacted in period *t* and city *c*	826	99	833	97	820	100
Avg_Rooms_{tc}	The average number of rooms of assets transacted in period *t* and city *c*	3.53	0.29	3.53	0.29	3.53	0.29
Avg_Age_{tc}	The average age (in years) of assets transacted in period *t* and city *c*	27.62	7.34	26.01	7.19	29.11	7.17
Avg_SES_{tc}	The average score on the socio-economic index of the statistical area where the asset is located	0.18	0.65	0.20	0.65	0.16	0.66
SD_Area_{tc}	The standard deviation of the area (in square feet) of assets transacted in period *t* and city *c*	271	54	276	55	267	54
SD_Rooms_{tc}	The standard deviation of the number of rooms of assets transacted in period *t* and city *c*	0.85	0.10	0.87	0.10	0.84	0.10
SD_Age_{tc}	The standard deviation of the age of assets transacted in period *t* and city *c*	13.31	3.17	12.92	3.24	13.68	3.07
SD_SES_{tc}	The standard deviation of the score on the socioeconomic index of the statistical area where the asset is located	0.47	0.19	0.47	0.19	0.47	0.19
SD_{tc}	Standard deviation of the residuals from the estimation of the price Eq. (1)	0.21	0.06	0.23	0.07	0.19	0.05
$P75-P25_{tc}$	Difference between the residuals in the 75th and the 25th percentiles	0.25	0.08	0.27	0.10	0.23	0.07

Notes: Table 2 presents summary statistics for the sample of city panel (all sample and sample stratified by pre- and post-information shock). Housing transaction data provided by the Israel Tax Authority; stock price data provided by the Tel Aviv Stock Exchange; all other data provided by the Israel Central Bureau of Statistics.

tion on the location score on the socioeconomic index (*SES*) is available from the Israel Central Bureau of Statistics.

4. Methodology

Consider the following estimated empirical model that examines the effect of the improved price information shock on price dispersion.

$$\ln(P_{itc}) = \beta_{0,c} + \vec{\beta}_{1,c}CHARACTERISTICS_{itc} + \vec{\beta}_{2,c}TFE_t + \varepsilon_{1itc} \text{ for all } c \quad (1)$$

and

$$SD_{tc} = \alpha_0 + \alpha_1 Treatment_t + \vec{\alpha}_2 CONTROLS1_{tc} + \vec{\alpha}_3 LFE_c + \varepsilon_{2tc}, \quad (2)$$

where Eq. (2) examines the effect of the information shock on price dispersion, and Eq. (1) is an auxiliary equation whose objective is to estimate the price dispersion to be substituted into Eq. (2), as further described below.

Eq. (1) is a hedonic price equation estimated for each city *c*. We use the estimation of Eq. (1) to generate the standard deviation of the residuals ε_{1itc} for every *t* and *c*, SD_{tc}, to be substituted on the left-hand side of Eq. (2). The dependent variable in Eq. (1), $\ln(P_{itc})$, is the log of the closing price of transaction *i* at time *t* in city *c*, and the independent variables in (1) include a series of asset characteristics, $CHARACTERISTICS_{itc}$, comprised of a vector of dummy variables $ROOMS\_Z$ ($Z=2,3,\ldots,6$), where $ROOMS\_Z_i=1$ if unit *i*'s number of rooms is equal to either *Z* or *Z*-0.5 and zero otherwise; *Area*, the floor area (in square feet); *Age*, the logarithm of the structure's age (in years); *SES*, the socioeconomic index score of the statistical area where the asset is located within the city[16]; and TFE_t is a vector of time-fixed effects. Also, β_0 is an estimated parameter, $\vec{\beta}_1$ and $\vec{\beta}_2$ are vectors of estimated parameters, and ε_{1itc} is a random disturbance term.[17]

The dependent variable in Eq. (2), SD_{tc}, is the standard deviation of ε_{1itc} that follows from Eq. (1), where subscripts *i, t*, and *c* denote transactions, quarters, and cities, respectively. The independent variables in Eq. (2) include $Treatment_t$, indicating post-information shock periods (a dummy variable that equals 1 for post-October 2010 periods and zero for

[16] A statistical area—the Israeli equivalent of a census tract—is the smallest geographic area examined by the Israel Central Bureau of Statistics (see more on this geographical unit in Section 6 below). The socioeconomic index (provided by the Israel Central Bureau of Statistics) may range from −3 to +3 and is generated by 16 indicators of the statistical area, clustered into 4 groups: standard of living, employment and welfare, schooling and education, and demography (see Israel Central Bureau of Statistics, 2013).

[17] While the log transformation in Eq. (1) reduces potential heteroskedasticity (see, among many others, Clemons et al., 2002), note that the validity of our test in Eq. (2) maintains even if heteroskedasticity exists in (1), as our focus in this auxiliary equation is to derive the standard deviation of ε_{1itc} rather than test for the significance of the coefficients.

pre-April 2010 periods) and a vector of control variables, $CONTROLS1_{tc}$, comprised of N_{tc}, the number of transactions at time t in city c, reflecting the amount of information that is generated by market depth; $\Delta\hat{P}_{tc}$, the 6-month (ending at t) rate of change in quality-adjusted housing prices in city c, controlling for the changes in the price level that may associate with price dispersion; $SD\_\hat{P}_{tc}$, the first difference in the 6-month (ending at t) moving standard deviation of quality-adjusted monthly housing prices in city c, controlling for the volatility in the time-series of the price that may affect the time t cross-sectional (across transacted units) quality-adjusted price dispersion (see derivation of $\Delta\hat{P}_{tc}$ and $SD\_\hat{P}_{tc}$ in the Appendix); SD_Stock_t, the first difference in the quarterly standard deviation of daily yields on the Tel Aviv 100 stock index, proxying the current level of uncertainty in the economy; and $Avg\_Attributes_{tc}$ and $SD\_Attributes_{tc}$, respective vectors of the average and standard deviation of dwelling attributes (across transacted dwellings at each couplet t and c), controlling for potential correlation between SD_{tc} and the distribution of dwelling unit attributes across time and space.[18] Finally, LFE_c on the right-hand side of (2) is a city fixed-effect indicator, and the parameters α_0 and α_1 are estimated coefficients, $\vec{\alpha}_2$ and $\vec{\alpha}_3$ are vectors of estimated coefficients, and ε_{2tc} is a random disturbance term.

The derivation of SD_{tc} in Eq. (1) and its substitution in the panel specification of Eq. (2) are designed to test the effect of improved information shock on price dispersion. We anticipate that the sudden availability of price information is followed by a decreased standard deviation of the residuals [from Eq. (1)], that is, that $\alpha_1 < 0$ in estimated Eq. (2).

In sum, based on the universe of all housing transactions, we estimate a hedonic price model in Eq. (1) for each city c [for a total of 57 estimations of Eq. (1); importantly, all outcomes reported below are robust to estimating Eq. (1) separately for all c and t while omitting the time-fixed effects—thereby allowing for time-varying coefficients in Eq. (1)]. Following this first step, we compute the standard deviation of the residuals in each city c for every quarter t to generate SD_{tc}, and then employ an unbalanced quarterly panel data of all cities over the period 2007–2013 (1264 observations in total)—based on which we estimate Eq. (2) to test the effect of the price information shock on subsequent price dispersion.

5. Results

Fig. 2A depicts the time series of the period t average (across cities) of ε_{2tc} over the period 2005Q1–2013Q4 that follows from the estimation of Eq. (2)—after excluding the variable *Treatment* from the estimation.[19] The figure shows a preliminary visual indication for the controlled effect of the information shock on decreased price dispersion. Whereas the time fixed-effects of the controlled SD_{tc} prior to the shock fluctuate around positive values, they turn negative in post-shock periods.

Table 3 presents the results of panel estimation that tests for the effect of the price information disclosure shock on the subsequent dispersion of quality-adjusted transaction closing prices. Column 1 presents the outcomes obtained from the estimation of Eq. (2) over the period 2007Q2–2010Q1 (12 quarters of pre-information shock) and 2011Q1–2013Q4 (12 quarters of post-information shock).[20] Empirical findings provide evidence to support an information effect on the dispersion of quality-adjusted prices. The coefficient on the *Treatment* variable is negative and significant at the 1% level. In particular, improved information shock associates with a decreased *SD* of 4.0% of property value. As the average standard deviation of the residuals in the period prior to the price disclosure equals 0.23 (see Table 2), this implies roughly an 18% decrease in price dispersion due to improved information shock.[21]

Column 2 in Table 3 presents the outcomes from re-estimating Eq. (2) over the periods 2009Q2–2010Q1 and 2011Q1–2011Q4 (that is, 4 quarters prior to and subsequent to the information shock, respectively). It follows that the short-term (4-quarter-subsequent) effect of the improved information shock is also salient and significant at the 1% level. The coefficient on the *Treatment* variable is equal to 2.7%, implying about a 12% decrease in price dispersion.

Further, we repeat the estimation of Eq. (2) for the 2007–2013 and 2009–2011 periods, substituting the *SD* measure of dispersion on the left-hand side of (2) with *P75-P25*, the difference between the residuals, ε_{1itc}, in the 75th and 25th percentiles (of the residual distribution) that follow from the estimation of Eq. (1) (summary statistics of *P75-P25* are presented in Table 2). This price dispersion measure is robust to outliers in the price observations. Results of this specification are presented in columns 3 and 4 of Table 3. It follows that estimation outcomes are robust to this specification. The improved information shock associates with about a 5.4% (4.4%) decrease, significant at the 1% level, in *P75-P25* for the 12- (4-) quarter pre- and post-treatment time frame. As the average value of *P75-P25* prior to the information shock is equal to 0.27, our outcome indicates about a 20% (17%) decrease in the price dispersion over the 12- (4-) quarter period subsequent to the information shock under this alternative measure.[22]

Results are further robust to: (*a*) changing the left-hand-side variable in Eq. (2), *P75-P25*, with either *P90-P10* or *Pmax-Pmin*, the difference between the residuals in the 90th and the 10th percentiles, or the difference between the maximum and minimum, respectively, of the residual distribution that follows from the estimation of Eq. (1); (*b*) the substi-

[18] Note that we test for possible endogeneity between SD_{tc} and N_{tc} in Eq. (2), where we use the number of construction starts and the rate of change in gross domestic product as instrumental variables for *N* (both are significant at the 1%-level in the reduced form equation for *N*). The Wu Hausman test, however, does not reject the hypothesis of no endogeneity between SD_{ct} and N_{ct} (p-value = 0.90). We thus estimate (2) under the no endogeneity assumption [yet all outcomes below are also robust to using 2SLS in the estimation of (2)]. Finally, to address potential endogeneity between SD_{tc} and either $\Delta\hat{P}_{tc}$ or $SD\_\hat{P}_{tc}$, we also re-estimate Eq. (2) based on lagged values ($\Delta\hat{P}_{t-1c}$ and $SD\_\hat{P}_{t-1c}$)—all results are robust to this alternative specification.

[19] That is, the line in Fig. 2A presents the period t average (across cities) of ε_{2tc} that follows from estimating the equation $SD_{tc} = \alpha_0 + \vec{\alpha}_2 CONTROLS1_{tc} + \vec{\alpha}_3 LFE_c + \varepsilon_{2tc}$. We also reject the hypothesis that the controlled SD_{tc} exhibits a negative trend prior to the information shock by estimating $SD_{tc} = \alpha_0 + \vec{\alpha}_2 CONTROLS1_{tc} + \vec{\alpha}_3 LFE_c + \alpha_4 t + \varepsilon_{2tc}$ over the pre-treatment period, where t represents quarters (over the period 2005Q1–2010Q1). The estimated coefficient on t is about zero ($\hat{\alpha}_4 = -0.0002$) and statistically insignificant. (Other outcomes from this test are not reported but are available on request).

[20] We use clustered standard errors in the estimation of Eq. (2) as the homoskedasticity assumption is rejected (χ^2 (57) = 1546; p-value < 0.0001), and we further reject the no-serial correlation assumption in (2) (Wooldridge test generates F-statistic (1,55) = 19.787; p-value < 0.001). For the 1-year pre- and post-treatment estimation of Eq. (2), we use robust standard errors, as we cannot reject the no-serial correlation assumption [Wooldridge test generates F-statistic (1,44) = 1.356; p-value = 0.250)]. Also, outcomes are robust to using weighted least-squares in the estimation of (2), where weights are determined by the total number of transactions in each city. The average R^2 coefficient from the estimations of the hedonic price Eq. (1) is equal to 0.80, ranging from 0.63 to 0.89 [outcomes from the estimations of auxiliary Eq. (1) are not reported but are available on request]. Summary statistics of SD_{tc}, the standard error of the residuals from the estimated price Eq. (1), are presented in Table 2. As can be seen, the average and standard deviation of SD_{tc} over the entire sample are about 0.21 and 0.06, respectively. Finally, as we use quarterly time-units, our post information shock estimation begins in 2011Q1. Outcomes, however, are robust to beginning the post-information shock estimation in 2010Q4.

[21] Note that the standard deviation of the residuals from Eq. (1), *SD*, is estimated in log of asset price. Hence, the residuals represent percentage of asset value.

[22] We also re-estimate Eq. (2) over the 12-quarter pre/post period, replacing the dummy variable *Treatment* on the right-hand side of (2) with three dummy variables indicating, respectively, one, two, and three years after the shock. The obtained coefficients are equal to −0.034, −0.048, and −0.054, respectively (all significant at the 1%-level)—indicating that the full impact of the shock is attained gradually (we thank the anonymous referee for pointing out this issue).

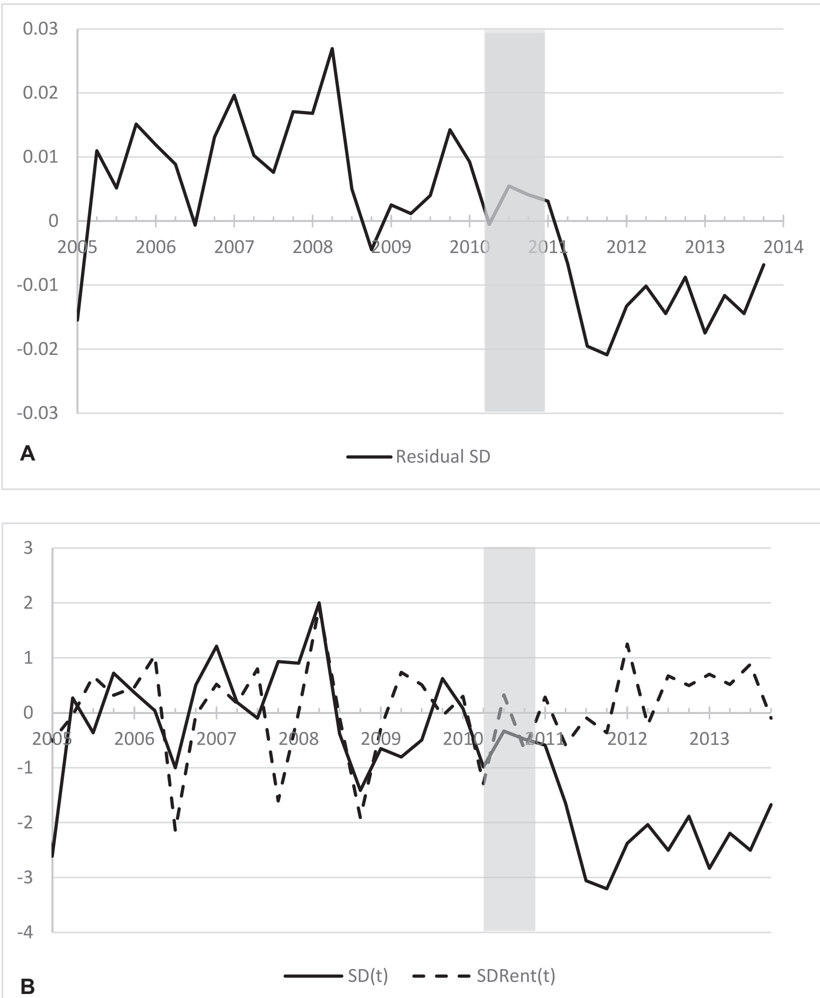

Fig. 2. (A) Period t Average (across Cities) of the Residuals ε_{2tc} that Follow from the Estimation of $SD_{tc} = \alpha_0 + \vec{\alpha}_2 CONTROLS1_{tc} + \vec{\alpha}_3 LFE_c + \varepsilon_{2tc}$ over the Period 2005Q1–2013Q4
Notes: The depicted values are the period t average (across cities) of ε_{2tc} that follows from the estimation of $SD_{tc} = \alpha_0 + \vec{\alpha}_2 CONTROLS1_{tc} + \vec{\alpha}_3 LFE_c + \varepsilon_{2tc}$. Pre- (post-) information shock periods are to the left (right) of the shaded area.
(B). Period t Average (across Cities) of the Residuals ε_{4tc}, Stratified by Sale and Rental Transactions, that Follow from the Estimation of $Y_{tc} = \mu_0 + \vec{\mu}_3 CONTROLS1_{tc} \times Sale_{tc} + \vec{\mu}_4 CONTROLS2_{tc} \times (1 - Sale_{tc}) + \vec{\mu}_5 LFE_c \times Sale_{tc} + \vec{\mu}_6 LFE_c \times (1 - Sale_{tc}) + \varepsilon_{4tc}$ over the period 2005Q1–2013Q4
Notes: The depicted values are the period t average (across cities) of ε_{4tc} (stratified by sale and rental transactions) that follows from the estimation of $Y_{tc} = \mu_0 + \vec{\mu}_3 CONTROLS1_{tc} \times Sale_{tc} + \vec{\mu}_4 CONTROLS2_{tc} \times (1 - Sale_{tc}) + \vec{\mu}_5 LFE_c \times Sale_{tc} + \vec{\mu}_6 LFE_c \times (1 - Sale_{tc}) + \varepsilon_{4tc}$. Pre- (post-) information shock periods are to the left (right) of the shaded area. For ease of inspection, SD_t and $SDRent_{t-1}$ are standardized (hence, y-axis is in units of standard deviation of SD_t and $SDRent_t$).

tution of the left-hand-side variable in (2) with its natural logarithm; (*c*) the omission of the city fixed-effect variables (LFE_c), the omission of the average attributes ($Avg\_Attributes$) and their standard deviation (*SD_Attributes*), and the omission of the vector *CONTROLS*1 on the right-hand side of (2); (*d*) the omission of all control variables, while maintaining only the city fixed-effects (LFE_c); (*e*) including only the 50% most active cities (transaction-wise) in the sample; (*f*) excluding the 50% most active cities from the sample; and (*g*) changing the condition for city/quarter participation in the estimation from a minimum of 20 transactions to a minimum of either 30, 40, or 50 transactions (results from these robustness tests are not reported but are available upon request).

Finally, the estimated coefficients of the control variables are as follows: Consistent with previous literature (see, e.g., Gatti and Kattuman, 2003; Eckard, 2004), price dispersion negatively correlates with price change, as a 1% increase in the quality-adjusted price associates with a decrease in *SD* equal to 0.05% of property value (significant at the 1% level and equivalent to about a 0.2% decrease in price dispersion). In addition, an increase in the 6-month (ending at *t*) time-series standard deviation of quality-adjusted housing prices, as well as an increase in the standard deviation of stock prices, associates with an increased *SD* (both significant at the 1% level).

6. Additional identification and robustness tests

In this section, we present a series of tests that assess the robustness of our findings to issues of test design and sampling.

6.1. Rental market as a control group

As the sale price information shock covered nationwide transactions, there is no market of housing sales that may serve as a pre/post control group for our investigation. Nevertheless, as housing sale and rental markets share common underlying economic forces, the local rental housing market immediately arises as a natural candidate to serve as a control group. In fact, up-to-date, micro-level rental housing price information has never been publicly disclosed in Israel. As it used to be with sale prices prior to April 2010 (before sale transaction information was first disclosed), the primary means of collecting rental price information is by observing asked prices either on designated private websites or in private listings managed by local real estate agents. We therefore use rental housing transactions as a control group for studying the effect of the information shock on sale price dispersion.[23] We hypothesize that, while information may spill over from the sale to rental market, as the information shock includes (excludes) sale (rent) prices, the decline in sale price dispersion is more salient than that of rental prices.

Based on a sample of about 82,700 rental housing transactions over the period 2005–2013 available to us from the Israel Central Bureau of Statistics (ICBS), we extend the pre/post estimation of the model in

[23] Commercial real estate transactions are another potential candidate for a control group. Unfortunately, however, there are no data available on this segment of the market.

Table 3
Regression results for the city-level panel estimation of Eq. (2).

Column Dependent variable	(1) *SD*	(2) *SD*	(3) *P75-P25*	(4) *P75-P25*
# of quarters prior and subsequent to the treatment	12 quarters	4 quarters	12 quarters	4 quarters
Constant	0.043	0.02	−0.051	−0.002
	(0.038)	(0.094)	(0.064)	(0.207)
$Treatment_t$	−0.040***	−0.027***	−0.054***	−0.044***
	(0.002)	(0.005)	(0.004)	(0.010)
N_{tc}	0.001	−0.001**	0.001	−0.001**
	(0.001)	(0.001)	(0.001)	(0.001)
$\Delta\hat{P}_{tc}$	−0.051***	−0.008	−0.110***	−0.111
	(0.015)	(0.038)	(0.025)	(0.104)
SD_P_{tc}	0.152***	0.105	0.237**	0.247
	(0.044)	(0.140)	(0.074)	(0.218)
SD_Stock_t	0.006***	−0.003	0.010***	−0.008
	(0.002)	(0.003)	(0.003)	(0.006)
$AVG\_Attributes_{ct}$	Included	Included	Included	Included
$SD\_Attributes_{ct}$	Included	Included	Included	Included
LFE (city fixed-effect)	Included	Included	Included	Included
# of Observations	1270	434	1270	434
# of Cities	57	57	57	57
R^2 (within cities)	0.371	0.233	0.261	0.184
Prob> F	0.000	0.000	0.000	0.000
Spatial unit	City	City	City	City
Temporal unit	Quarter	Quarter	Quarter	Quarter

Notes: Table 3 presents results of GLS with fixed-effects estimation of Eq. (2) with clustered standard errors for the 12 quarters estimation and robust standard error for the 4 quarters estimations. Columns 1 and 3 (2 and 4) present estimation results for the period that includes 12 (4) quarters prior to and subsequent to the information shock. Columns 3 and 4 further present estimation results when *P*75-*P*25 replaces *SD* on the left-hand side of Eq. (2). Robust standard errors are provided in parentheses. One, two, and three asterisks represent significance at the 10%, 5%, and 1% levels, respectively.

Table 4
List of variables, description, and summary statistics of rental transactions.

Variable	Sample	All		Pre-shock		Post-shock	
	Description	Avg.	Std.	Avg.	Std.	Avg.	Std.
P	Rental price (in dollars)	742.2	313.6	642.2	251.5	856.7	340.0
Rooms_2	Dummy variable that equals 1 if number of rooms is equal to 1.5 or 2; zero otherwise	0.15	0.36	0.15	0.36	0.15	0.35
Rooms_3	Dummy variable that equals 1 if number of rooms is equal to 2.5 or 3; zero otherwise	0.45	0.50	0.45	0.50	0.44	0.50
Rooms_4	Dummy variable that equals 1 if number of rooms is equal to 3.5 or 4; zero otherwise	0.29	0.45	0.28	0.45	0.29	0.46
Rooms_5	Dummy variable that equals 1 if number of rooms is equal to 4.5 or 5; zero otherwise	0.08	0.27	0.08	0.28	0.08	0.27
Rooms_6	Dummy variable that equals 1 if number of rooms is equal to 5.5 or 6; zero otherwise	0.01	0.12	0.02	0.12	0.01	0.11
Water	Dummy variable that equals 1 if landlord pays for water; zero otherwise	0.02	0.15	0.02	0.14	0.02	0.15
Tax	Dummy variable that equals 1 if landlord pays related municipal taxes; zero otherwise	0.03	0.17	0.02	0.16	0.03	0.18
Maintenance	Dummy variable that equals 1 if landlord pays for maintenance; zero otherwise	0.02	0.15	0.02	0.14	0.03	0.16
Electricity	Dummy variable that equals 1 if landlord pays for electricity; zero otherwise	0.01	0.10	0.01	0.10	0.01	0.10
Duration	Duration of lease (in months)	11.00	4.92	10.58	3.78	11.69	6.05
SES	Score on a socioeconomic index of the statistical area where the property is located	11.31	3.45	11.34	3.48	11.28	3.41
Continue	Dummy variable that equals 1 if the current lease continues a previous lease contract between the parties	0.66	0.47	0.66	0.47	0.67	0.47
Adjusted	Dummy variable that equals 1 if the lease is price-level adjusted	0.02	0.13	0.01	0.12	0.02	0.14
Dollar	Dummy variable that equals 1 if the lease is adjusted to the shekel-dollar exchange rate	0.27	0.45	0.52	0.50	0.02	0.15
Downpayment	Number of pre-paid monthly payments	1.52	1.47	1.65	1.63	1.38	1.27

Notes: Table 4 presents summary statistics of the sample of rental transactions (all sample and sample stratified by pre- and post-information shock). The socioeconomic index (*SES*) for rental transactions provided by ICBS is equivalent to the socioeconomic index of the sale transactions; however, it is translated by ICBS to a scale ranging from 1 to 20. The rent amount was translated from NIS to US dollars under 1USD = 3.77NIS.

Eqs. (1) and (2) to include rental market transactions.[24] In total, 12 major cities and 16 districts (each district includes a number of towns) in Israel are included in the rental transaction sample—altogether representing all major rental markets in Israel (see Table 4 for the list of variables and summary statistics of rental transactions—all sample and sample stratified by pre- and post-information shock). The variables in the sample include asset and neighborhood characteristics (including number of rooms and socioeconomic index of the statistical area where the asset is located), as well as lease contract attributes (including the duration of the lease, whether the price is adjusted to inflation, whether a down payment is required, the party responsible for paying current water, tax, and maintenance expenses, and whether the lease is new or continued).

We use the sale and rental transaction datasets to estimate Eq. (1) (see Section 4 above), together with the following:

$$\ln\left(P^{Rent}_{itc}\right) = \pi_{0,c} + \vec{\pi}_{1,c}CHARACTERISTICS2_{itc} + \vec{\pi}_{2,c}CONTRACT_{itc} + \vec{\pi}_{3,c}TFE_t + \varepsilon_{3itc} \text{ for all } c \tag{1a}$$

[24] The total raw sample includes 85,251 rental housing observations—about 450 to 1500 per month—collected and used by ICBS to periodically estimate the change in the housing user cost index. We require a minimum of 20 observations per city/quarter to be included in the sample. Yet outcomes (which are not reported but are available on request) are robust to requiring a minimum of either 30 or 40 observations per city/quarter.

Table 5
List of city-level panel rental variables and summary statistics.

Sample	All		Pre-shock		Post-shock	
Variable	Avg.	Std.	Avg.	Std.	Avg.	Std.
$Treatment_t$	0.366	0.482	0	0	1	0
Avg_Rooms	3.308	0.329	3.31	0.35	3.31	0.30
Avg_SES	10.822	2.116	10.80	2.15	10.84	2.08
Avg_Water	0.0	0.0	0.02	0.04	0.03	0.05
Avg_Tax	0.033	0.049	0.03	0.04	0.04	0.06
Avg_Maintenance	0.020	0.032	0.02	0.04	0.02	0.03
Avg_Electricity	0.010	0.023	0.01	0.03	0.01	0.02
Avg_Duration	10.861	1.334	10.57	1.01	11.56	1.58
Avg_Continue	0.645	0.137	0.64	0.15	0.65	0.12
Avg_Adjusted	0.014	0.026	0.01	0.03	0.02	0.02
Avg_Dollar	0.307	0.354	0.51	0.34	0.02	0.03
Avg_Downpayment	1.554	0.395	1.67	0.42	1.37	0.29
SD_Rooms	0.892	0.195	0.89	0.21	0.89	0.17
SD_SE	2.385	0.940	2.41	0.94	2.34	0.96
SD_Duration	4.136	2.013	3.52	1.31	5.21	2.62
SD_Downpayment	1.329	0.733	1.45	0.76	1.13	0.66
SD_Rent	0.200	0.047	0.20	0.05	0.20	0.04
P75-P25_Rent	0.266	0.085	0.27	0.09	0.27	0.08

Notes: Table 5 presents summary statistics for the sample of city-level rental panel (all sample and sample stratified by pre- and post-information shock).

and

$$Y_{tc} = \mu_0 + \mu_1 Treatment_t + \mu_2 Treatment_t \times Sale_{tc} + \vec{\mu}_3 CONTROLS1_{tc} \times Sale_{tc} + \vec{\mu}_4 CONTROLS2_{tc} \times (1 - Sale_{tc}) + \vec{\mu}_5 LFE_c \times Sale_{tc} + \vec{\mu}_6 LFE_c \times \left(1 - Sale_{tc}\right) + \varepsilon_{4tc}, \quad (3)$$

where Eqs. (1) and (1a) serve as auxiliary hedonic price equations whose objective is to estimate the quality-adjusted sale and rental price dispersion, respectively, and Eq. (3) is designed to directly examine the difference in the effect of the information shock on sale and rental price dispersion, as further explained below.

Eq. (1a) is a rental version of the hedonic sale price Eq. (1). The dependent variable in Eq. (1a), $\ln(P^{Rent}_{itc})$, is the log of rental price of housing unit *i* at time *t* in city (or district) *c*, and the independent variables in (1a) include asset characteristics, *CHARACTERISTICS*2, comprised of a vector of dummy variables, *ROOMS_X* ($X = 2,3,\ldots,6$), where $ROOMS\_X_i = 1$ if unit *i*'s number of rooms is equal to *X* (or *X*–0.5) and zero otherwise; and *SES*, the score on a socioeconomic index of the statistical area where property *i* is located. Also, *TFE* is a vector of time-fixed effects and *CONTRACT* is a vector of rental contract attributes, comprised of *Downpayment*, the required number of pre-paid months; *Duration*, the duration of the lease; and a series of dummy variables including *Continue*, indicating whether the current lease continues a previous lease contract between the parties; *Adjusted*, indicating whether the lease is price-level adjusted; *Dollar*, indicating whether the lease is adjusted to the shekel-dollar exchange rate; and *Water, Electricity, Maintenance*, and *Tax*, indicating whether the landlord pays for water, electricity, maintenance, and taxes, respectively (otherwise those are paid by the tenant). Finally, $\pi_{0,\,c}$ is an estimated parameter, $\vec{\pi}_{1,c}$–$\vec{\pi}_{3,c}$ are vectors of estimated parameters, and ε_{3itc} is a random disturbance term (summary statistics of the panel rental variables are presented in Table 5—for the entire sample as well as sample stratified by pre- and post-information shock).

We estimate equations (1) (based on the sale transactions) and (1a) (based on the rental transactions) over the period 2007–2013 to generate an unbalanced panel of the standard deviation of the residuals ε_{1itc} and ε_{3itc}, respectively, for every *t* and *c* (denoted by SD_{tc} and $SDRent_{tc}$, respectively). We then standardize SD_{tc} and $SDRent_{tc}$ (subtracting the pre-shock mean and dividing by the pre-shock standard deviation of each series) so that the series appear on the same scale, and substitute them for Y_{tc} on the left-hand side of Eq. (3) to estimate the effect of the information shock on sale and rental price dispersion. The independent variables in Eq. (3) include $Treatment_t$, indicating post-information shock periods (a dummy variable that equals 1 for the quarters subsequent to October 2010 and zero for quarters prior to April 2010); and an interaction term, $Treatment_t \times Sale_{tc}$, where $Sale_{tc}$, is a dummy variable that equals 1 for sale transactions and zero for rental transactions. In addition, the right-hand side of Eq. (3) includes two vectors of control variables: $CONTROLS1_{tc}$ is a vector of controls [that appeared on the right-hand side of Eq. (2)], controlling for the sale transactions (thus interacted with the dummy variable $Sale_{tc}$) and $CONTROLS2_{tc}$, an equivalent vector of controls for rental transactions [thus interacted with $(1 - Sale_{tc})$], comprised of N^{Rent}_{tc}, the sample number of rental transactions at time *t* in city *c*; $\Delta\hat{P}^{Rent}_{tc}$, the 6-month (ending at *t*) rate of change in quality-adjusted rental prices in city *c*; $Avg\_Rooms^{Rent}_{tc}$ and $SD\_Rooms^{Rent}_{tc}$, respective vectors of the average and standard deviation (across rental units at each couplet *t* and *c*) of the number of rooms; $Avg\_SE^{Rent}_{tc}$ and $SD\_SE^{Rent}_{tc}$, the average and standard deviation (across rental units at each couplet *t* and *c*) of the score on the socioeconomic index; and $Avg\_Attributes^{Rent}_{tc}$ and $SD\_Attributes^{Rent}_{tc}$, respective vectors of the average and standard deviation of rental unit physical and contract attributes (across units at each couplet *t* and *c*), controlling for potential correlation between $SDRent_{tc}$ and the distribution of physical and contract attributes across time and space. Finally, $LFE_c \times Sale_{tc}$ and $LFE_c \times (1 - Sale_{tc})$, on the right-hand side of (3) are city fixed-effects, separately indicating rental and sale transactions, $\mu_0 - \mu_2$ and $\vec{\mu}_3 - \vec{\mu}_6$ are estimated parameters and vectors of parameters, respectively, and ε_{4tc} is a random disturbance term.[25]

Fig. 2B depicts the time series of the period *t* average (across cities) of ε_{4tc} [that follows from the estimation of Eq. (3) after excluding the variables *Treatment* and *Treatment* × *Sale* from the right-hand side of the equation]—stratified by sale and rental observations over the period 2005Q1–2013Q4.[26] The figure provides a preliminary visual indication for both the parallel trends of the sale and rental price dispersion in the period preceding the information shock (Pearson correlation between the two series over the period 2005–2010Q1 is equal to 0.5, significant at the 5% level) and the diverse effect on the dispersion of sale and rental prices following the shock. As compared to the downtrend experienced by sale price dispersion in the period succeeding the information shock, post-shock rental price dispersion continued to fluctuate around roughly the same pre-shock level.

Table 6 presents the results of the panel estimation of Eq. (3) that tests for the diverse effect of the price information shock on quality-adjusted sale and rental price dispersion. Column 1 presents the outcomes obtained for the period 2007Q2–2010Q1 (12 pre-shock quarters) and 2011Q1–2013Q4 (12 post-shock quarters). Empirical findings provide evidence in support of a salient (no) effect of the sale price information shock on the dispersion of quality-adjusted sale (rent) prices. The coefficients on *Treatment* and *Treatment* × *Sale* are equal to 0.08 (insignificant) and –0.68 (significant at the 1% level), respectively, implying that the dispersion of sale prices decreases by about 0.6 standard deviations of *SD*. As the pre-shock standard deviation of *SD* is equal to 0.07, the latter implies a decreased price dispersion equal to 4% (0.6 × 0.07) of the average sale price, which is translated to about a 18% decrease in sale price dispersion due to improved information shock (recall that the pre-shock average of *SD* is equal to 0.23). At the same time, however, rental price dispersion is largely insensitive to the information shock.

Column 2 in Table 6 presents the outcomes from re-estimating Eq. (3) over the periods 2009Q2–2010Q1 and 2011Q1–2011Q4 (i.e., 4 quarters prior and subsequent to the information shock). Columns 3

[25] Note that we omit the variable $Sale_{tc}$ on the right-hand side of (3), as it is collinear with $LFE_c \times Sale_{tc}$ and $LFE_c \times (1 - Sale_{tc})$.

[26] Our evidence shows that prior to the information shock, while SD_{tc} and $SDRent_{tc}$ maintain similar dynamics, the time-variance of $SDRent_{tc}$ often precedes that of SD_{tc} by one period (quarter). For ease of visual inspection, Fig. 2B thus presents the period *t* and t−1 average (across cities) of ε_{4tc} stratified by sale and rental observations, respectively.

Table 6
Regression results for the sale and rental market estimation of Eq. (3).

Column	(1)	(2)	(3)	(4)
Dependent variable	*SD*	*SD*	*P75-P25*	*P75-P25*
# of quarters prior and subsequent to the treatment	12 quarters	4 quarters	12 quarters	4 quarters
Constant	-3.317***	-3.597***	-3.169***	-2.944*
	(0.423)	(0.985)	(0.503)	(1.551)
$Treatment_t$	0.077	0.071	0.036	0.073
	(0.049)	(0.049)	(0.058)	(0.109)
$Treatment_t \times Sale_{tc}$	−0.675***	−0.472***	−0.569***	−0.508***
	(0.058)	(0.091)	(0.069)	(0.144)
$CONTROLS1_{tc}$	Included	Included	Included	Included
$CONTROLS2_{tc}$	Included	Included	Included	Included
LFExSale	Included	Included	Included	Included
LFE	Included	Included	Included	Included
# of Observations	1899	648	1899	648
# of Groups	85	85	85	85
R^2 (within cities)	0.404	0.353	0.252	0.254
Prob> Chi2	0.000	0.000	0.000	0.000
Spatial unit	City	City	City	City
Temporal unit	Quarter	Quarter	Quarter	Quarter

Notes: Table 6 presents results of GLS with fixed-effects estimation of Eq. (3). We use robust standard errors, as we do not reject the no-serial correlation assumption (test results are not reported but are available upon request). Columns 1 and 3 (2 and 4) present estimation results for the period that includes 12 (4) quarters prior to and subsequent to the information shock. Columns 3 and 4 further present estimation results when *P75-P25* respectively replaces *SD* and *SDRent* on the left-hand side of Eq. (3). Robust standard errors are provided in parentheses. One, two, and three asterisks represent significance at the 10%, 5%, and 1% levels, respectively.

and 4 in the table present the outcomes from re-estimating the equation after substituting Y on the left-hand side of (3) with *P*75-*P*25, the difference between the residuals in the 75th and the 25th percentiles (of the residual distribution) that follow from the estimation of Eqs. (1) and (1a)—for the 2007–2013 and 2009–2011 periods, respectively. It follows from columns 2–4 that the salient (no) effect on sale (rental) price dispersion is robust to these specifications.[27]

6.2. Smaller geographical areas and pseudo control groups

The estimation of Eq. (2) reported above shows that improved information shock associates with a considerable decrease in the dispersion of subsequent sale transaction prices. This outcome is based on a panel of quarterly sale observations in the 57 most active cities. We now augment those findings on the correlation between the price information shock and subsequent price dispersion by focusing on smaller geographical areas. The Israel Central Bureau of Statistics (ICBS) divides all municipalities in Israel hosting no fewer than 10,000 residents into geographical units referred to as statistical areas (the smallest sampling geographical unit employed by ICBS), which are equivalent to census tracts in the United States. Each statistical area includes about 3000–5000 residents, and, as with census tracts, the division into statistical areas accounts for aspects of homogeneity with respect to population characteristics, economic status, and living conditions (see ICBS, 2013).[28]

The analytic gain from using these smaller geographical units comes, however, with a decreased number of transactions per location per period. We thus extend the time-unit of the statistical area panel to one year. Altogether, our panel thus includes all housing transactions in 829 statistical areas (a total of 174,611 observations) over the 2007–2013 period. Table 7 presents summary statistics of this sample (all sample and sample stratified by pre- and post-information shock). As indicated in the table, the average dwelling unit across statistical areas is a 3.4-room, 843-square-foot condominium apartment located in a 30-year-old structure.

We re-estimate Eqs. (1) and (2) above based on a statistical area-level sample, where subscripts t and c are substituted by subscripts τ and s—representing annual time periods and statistical areas, respectively. We should note that the controls on the right-hand side of Eq. (2) now differ, however, from those estimated for the city-level sample in two ways. First, *SES* (the socioeconomic index score of the statistical area where the asset is located) is omitted, as it is only available by statistical areas and thus does not vary within the statistical area. In addition, we supplement the controls on the right-hand side of (2) with $\Delta\hat{P}_{\tau s}$, the 12-month (ending at τ) rate of change in quality-adjusted housing prices in the city to which statistical area s belongs, controlling for the changes in the price level that may associate with price dispersion; and $SD\_\hat{P}_{\tau s}$, the 12-month (ending at τ) standard deviation of quality-adjusted housing prices in the city to which statistical area s belongs, controlling for the volatility in the time-series of the price that may affect the time t cross-sectional (across transacted units) quality-adjusted price dispersion (see derivation of $\Delta\hat{P}_{\tau s}$ and $SD\_\hat{P}_{\tau s}$ in the Appendix).

Outcomes from the estimation of Eq. (2) are robust to the statistical area sample specification [results from the 829 estimations of (1) and the estimation of (2) based on the statistical area sample are not reported but are available upon request]. Specifically, it follows that the coefficient on the *Treatment* variable is equal to −0.024 (significant at the 1% level). Given that the average (across statistical areas) standard deviation of the residuals in the period prior to the price disclosure is 0.18, it follows that adjusted-price dispersion decreases by about 13%, *ceteris paribus*, subsequent to the price information shock. Further, the decreased price dispersion effect maintains when we limit the examined period to one year prior to and subsequent to the information shock (price dispersion decreases by about 13% as well). Finally, the outcome on the information shock effect is insensitive to substituting the price dispersion measure *SD* with *P75-P25*, the difference between the residual in the 75th and the 25th percentiles that follow from the estimation of Eq. (1) based on the statistical area sample.[29]

[27] Results are further robust to (*a*) the omission of $CONTROLS1_{tc}$ and $CONTROLS2_{tc}$ on the right-hand side of (3); (*b*) including only cities that are available in both the sale and rent datasets; (*c*) replacing the standardized series of *SD* and *SDRent* with the non-standardized series; and (*d*) replacing $LFE_c \times Sale_{tc}$ and $LFE_c \times (1 - Sale_{tc})$ with $Sale_{tc}$.

[28] For example, the three largest cities in Israel—Jerusalem, Tel Aviv, and Haifa—include 181, 164, and 91 statistical areas, respectively.

[29] We use clustered standard errors (robust standard errors) in the estimation of (2) based on statistical areas for the 3- (1-) year before/after periods as we

Table 7
List of statistical area-level panel variables, description, and summary statistics.

Variable	Sample	All		Pre-shock		Post-shock	
	Description	Avg.	Std.	Avg.	Std.	Avg.	Std.
$Treatment_{\tau}$	Dummy variable that equals 1 for periods subsequent to information shock (i.e., subsequent to the end of 2010)	0.50	0.50	0	0	1	0
$SD_{\tau s}$	The standard deviation of the residuals in price Eq. (1)	0.169	0.059	0.18	0.06	0.15	0.05
$P75-P25_{\tau s}$	Difference between the residuals in the 75th and 25th percentiles	0.204	0.082	0.23	0.09	0.18	0.07
$N_{\tau s}$	The number of transactions in period τ and statistical area s	40.5	17.0	40.94	16.46	39.98	17.51
$Avg\_Area_{\tau s}$	The average area (in square feet) of assets transacted in period τ and statistical area s	843	171	844.7	169.1	841.5	173.7
$Avg\_Rooms_{\tau s}$	The average number of rooms of assets transacted in period τ and statistical area s	3.41	0.48	3.40	0.49	3.41	0.48
$Avg\_Age_{\tau s}$	The average age (in years) of assets transacted in period τ and statistical area s	30.9	11.9	29.41	11.57	32.35	11.94
$SD\_Area_{\tau s}$	The standard deviation of the area (in square feet) of assets transacted in period τ and statistical area s	233	80	235.8	79.2	230.3	81.6
$SD\_Rooms_{\tau s}$	The standard deviation of the number of rooms of assets transacted in period τ and statistical area s	0.77	0.16	0.78	0.16	0.76	0.16
$SD\_Age_{\tau s}$	The standard deviation of the age of assets transacted in period τ and statistical area s	9.42	4.67	9.25	4.48	9.59	4.86
$\Delta\hat{P}_{\tau s}$	the 12-month (ending at τ) rate of change in quality-adjusted housing prices in the statistical area s	0.081	0.088	0.10	0.10	0.06	0.07
$SD\_\hat{P}_{\tau s}$	the 12-month (ending at τ) standard deviation of quality-adjusted housing prices in the city to which statistical area s belongs	0.040	0.021	0.05	0.02	0.03	0.01
SD_Stock_{τ}	1-year moving standard deviation of daily yields of the Tel Aviv 100 stock index	0.012	0.601	1.60	0.44	0.97	0.35
SES_s	The score of statistical area s on the socio-economic index	0.26	0.81	0.29	0.81	0.23	0.81

Notes: Table 7 presents summary statistics for the sample of statistical area panel (all sample and sample stratified by pre- and post-information shock). Transaction price data provided by the Israel Tax Authority; stock price data provided by the Tel Aviv Stock Exchange; all other data provided by the Israel Central Bureau of Statistics.

As a further robustness test of the information shock effect in the absence of "non-treated" housing markets, we stratify the sample by the level of statistical area price dispersion (*SD*) prior to the information shock. We hypothesize that if in certain sub-markets (statistical areas) some price information was already present (although not through formal information channels) prior to the formal price shock, then these sub-markets were to be less affected by the shock. In other words, sub-markets wherein price dispersion was *ex ante* relatively low were to be proportionally less impacted by the information shock.

Specifically, we stratify the sample by quartiles of the level of price dispersion (*SD*) prior to the information shock and re-estimate the following variation of Eq. (2) based on all observations in those statistical areas in the lowest (first) and highest (fourth) quartiles of *SD*:

$$SD_{\tau s} = \psi_0 + \psi_1 Treatment_{\tau} + \psi_2 High_s + \psi_3 High_s \times Treatment_{\tau} + \vec{\psi}_4 CONTROLS1_{\tau s} + \varepsilon_{5\tau s}, \tag{2a}$$

where, compared to (2), in the estimated Eq. (2a), we supplement two right-hand-side terms: $High_s$, a dummy variable denoting the high *SD* group, which equals one for a statistical area whose pre-treatment level of *SD* is in the fourth quartile and zero if the level of *SD* is in the first *SD* quartile; and $High_s \times Treatment_{\tau}$, an interaction term of $High_s$ and $Treatment_{\tau}$. Also, $\psi_0 - \psi_3$ and $\vec{\psi}_4$ are estimated parameters and vector of parameters, respectively; $\varepsilon_{5\tau s}$ is a random disturbance term; and all other variables are as described above. Provided that the low *SD* group is relatively less affected by the information shock (treatment), we expect that ψ_3 is significantly negative.[30]

Columns 1 and 2 in Table 8 present the outcomes from the estimation of Eq. (2a) based on the sub-sample of those statistical areas that appear in the first and fourth quartiles of the distribution of *SD*. As anticipated, the findings show that those statistical areas, which *ex ante* exhibit lower *SD*, are proportionally less affected by the treatment. In particular, focusing on ψ_1 and ψ_3, it follows from column 1 that the decrease in price dispersion over the three years after the treatment in statistical areas at the top (bottom) *SD* quartile is equal to 0.053 (0.006) (significant at the 1%-level). Given that the average *SD* in the period prior to the price disclosure is 0.26 (0.13) in the fourth (first) *SD* quartiles, it follows that adjusted-price dispersion decreases by about 20% (5%), *ceteris paribus*, subsequent to the price information shock. Column 2 in Table 8 shows that these results are robust to limiting the examined period to one year prior to and subsequent to the information shock.

Finally, as an additional robustness check, we stratify the sample by the rate of new housing sold in the statistical area in every period. Intuitively, in submarkets with a greater share of new units sold—whose price is published and likely available to market players—some price information may have been present prior to the information shock. One should therefore expect that sub-markets with a greater relative share of new housing sold were to be less impacted by the information shock.

We thus re-estimate Eq. (2a), where we respectively replace $High_s$ and $High_s \times Treatment_{\tau}$ with $Share\_New_{\tau s}$, a variable denoting the relative share of new housing sold in the statistical area s in period τ and $Share\_New_{\tau s} \times Treatment_{\tau}$, an interaction term of $Share\_New_{\tau s}$ and $Treatment_{\tau}$ (the mean and standard deviation of $Share\_New_{\tau s}$ are equal to 0.08 and 0.15, respectively). All other variables and parameters are as described in (2a). We expect that due to the price information that is provided by new housing sold, the share of new housing sold in a statistical area is associated with (*a*) a lower average price dispersion (i.e., $\psi_2 < 0$); and (*b*) a more moderate effect of the information shock (i.e., $\psi_3 > 0$).

Columns 3 and 4 in Table 8 present the outcomes from the estimation of Eq. (2a) based on the sample of statistical areas with $Share\_New_{\tau s}$ replacing $High_s$. It follows from column 3 that for the 3-year before/after

reject the homoskedasticity and (do not reject) the no-serial correlation assumptions (test results are not reported but are available upon request). Also, results are robust to: (*a*) WLS procedure, where each statistical area is weighted by the number of transactions; (*b*) the substitution of the dependent variable with its natural logarithm and with either *P*90-*P*10 or *Pmax-Pmin*, i.e., the difference between the 90th percentile and the 10th percentile, or the difference between the maximum and the minimum of the residuals from Eq. (1); and (*c*) limiting the sample to the three largest cities in Israel (Jerusalem, Tel Aviv, and Haifa). Summary statistics of the variables used in the statistical area panel estimation are presented in Table 7. The average and standard deviation of $SD_{\tau s}$ are 0.17 and 0.06, respectively, and of $P75-P25_{\tau s}$ are 0.20 and 0.08, respectively. Finally, the average R^2 of the 829 statistical area-level estimations of Eq. (1) is 0.78.

[30] As $High_s$ is specific to a statistical area s, we omit the vector of statistical area fixed-effects (LFE_s) from the estimation of (2a).

Table 8
Regression results for the statistical area-level estimation of Eq. (2a).

Column	(1)	(2)	(3)	(4)
Dependent variable	*SD*	*SD*	*SD*	*SD*
# of years prior to and subsequent to the treatment	3 years	1 year	3 years	1 year
Constant	0.090***	0.088***	0.140***	0.126**
	(0.016)	(0.025)	(0.019)	(0.052)
$Treatment_\tau$	−0.006***	0.006	−0.026***	−0.026***
	(0.002)	(0.005)	(0.002)	(0.004)
$High_s$	0.115***	0.115***		
	(0.003)	(0.005)		
$High_s \times Treatment_\tau$	−0.047***	−0.042***		
	(0.003)	(0.005)		
$Share\_New_{\tau s}$			−0.018**	0.014
			(0.007)	(0.016)
$Share\_New_{\tau s} \times Treatment_\tau$			0.028***	0.016
			(0.008)	(0.016)
$N_{\tau-1}$	0.000	0.000	−0.000	−0.000
	(0.000)	(0.000)	(0.000)	(0.000)
$SD\_P_{\tau s}$	0.254***	0.231**	0.167***	0.073
	(0.053)	(0.091)	(0.039)	(0.096)
$\Delta P_{\tau s}$	−0.018	−0.01	−0.049***	−0.012
	(0.011)	(0.022)	(0.007)	(0.026)
SD_Stock_τ	0.011***		0.016***	
	(0.001)		(0.003)	
$AVG\_Attributes_{\tau s}$	Included	Included	Included	Included
$SD\_Attributes_{\tau s}$	Included	Included	Included	Included
Number of observations	2185	717	4316	1391
Number of statistical areas	433	416	829	794
R^2 (within statistical areas)	0.35	0.24	0.24	0.11
Prob$>\chi^2$	0	0	0	0
Spatial unit	Statistical Area	Statistical Area	Statistical Area	Statistical Area
Temporal unit	Year	Year	Year	Year

Notes: Columns 1 and 2 in Table 8 present the outcomes from the estimation of Eq. (2a) based on the sub-sample of statistical areas that appear in the first and fourth quartiles of the distribution of *SD*. Columns 3 and 4 in Table 8 present the outcomes from the estimation of Eq. (2a) after stratifying the statistical area sample by the rate of new housing sold in every period. We use clustered standard errors (robust standard errors) in the estimation of (2a) for the 3- (1-) year before/after periods as we reject the homoskedasticity and (do not reject) the no-serial correlation assumptions. Standard errors are provided in parentheses. Two and three asterisks represent significance at the 5% and 1% levels, respectively.

test, the share of new housing sold is associated with (*a*) an average decrease in price dispersion (coefficient on $Share\_New_{\tau s}$ is equal to −0.018; significant at the 5%-level); and (*b*) a decreased effect of the information shock (coefficient on $Share\_New_{\tau s} \times Treatment_\tau$ is equal to 0.028; significant at the 1%-level). Specifically, it follows from the estimation that a 10% increase in the share of new housing sold in the statistical area both decreases the price dispersion prior to the information shock by about 1% and moderates the effect of the information shock on price dispersion by about 10%. Finally, it follows from column 4 in Table 8 that while the information shock effect maintains for the 1-year before/after test, the coefficients pertaining to new housing sold are insignificant.

6.3. Can unobservable variables explain our outcomes? Assessing the degree of omitted variables bias

To further gauge the robustness of our result on the decreased price dispersion that follows the improved information shock, we conduct a test, originally introduced by Altonji et al. (2005), assessing the relative amount of selection on potential unobservable variables in Eq. (2) that is required to explain the effect that we attribute to the information shock.[31]

Table 9 shows the results of the amount of selection on the unobservables relative to selection on the observables that is required in order to ascribe the entire information shock effect to selection bias, for various model specifications. It follows that the normalized shift in the distribution of the unobservables should be about 3.2 (8.7) greater than the normalized shift in the distribution of the *Treatment* (the observed information shock) to nullify the entire 3- (1-) year pre/post effect of the shock. Moreover, even under the assumption that selection on unobservables is *equal* to that on observables, the negative effect of the information shock on the price dispersion would maintain a statistical significance at the 1% level (see *t*-values in the fourth row of Table 9). In fact, the negative response of the price dispersion to the information shock would maintain significance at the 5%-level even if the shift in the distribution of the unobservables were about 2.8 (6.4) as great as that of the observables for the 3- (1-) year pre/post time-frame (see the fifth row of Table 9). Finally, it follows from columns 3 and 4 of Table 9 that all outcomes are robust to substituting *SD* on the left-hand side of (2) with *P*75-*P*25 for both the 3- and 1-year pre/post estimations.[32]

[31] Essentially, in the spirit of Eq. (2) above, Altonji et al. (2005) require that $\frac{E(\varepsilon_{2tc}|Treatment=1)-E(\varepsilon_{2tc}|Treatment=0)}{Var(\varepsilon_{2tc})} = \frac{E(\tilde{\alpha}_2 CONTROLS1_{tc}+\tilde{\alpha}_3 LFE_c|Treatment=1)-E(\tilde{\alpha}_2 CONTROLS1_{tc}+\tilde{\alpha}_3 LFE_c|Treatment=0)}{Var(\tilde{\alpha}_2 CONTROLS1_{tc}+\tilde{\alpha}_3 LFE_c)}$, i.e., that the relationship between *Treatment* and the mean of the distribution of the unobservables that generate our result is equal to the relationship between *Treatment* and the mean of the observables (both relationships adjusted by the respective variance). Their method estimates how large the ratio on the left-hand side of this condition would have to be relative to that on the right-hand side so as to account for the entire estimate of α_1 in Eq. (2)—under the null hypothesis of no information shock effect. See more on the derivation of this test in Altonji et al. (2005).

[32] Outcomes of our Altonji et al. (2005) test are further robust under the statistical area-level estimation of Eqs. (1) and (2).

Table 9
The amount of selection on the unobservables relative to selection on the observables required to ascribe the entire information shock effect to selection bias.

Column Dependent variable	(1) *SD*	(2) *SD*	(3) *P75-P25*	(4) *P75-P25*
Selection ratio for which the *Treatment* effect will be nullified	3.2	8.7	2.3	7.5
t-value in case of a 1:1 ratio	−11.5	−6.6	−7.7	−6.1
Maximum selection ratio for which the *Treatment* effect will maintain significance at the 5%-level	2.8	6.4	2.0	5.4
# of quarters prior and subsequent to the treatment	12 quarters	4 quarters	12 quarters	4 quarters
Spatial unit	City	City	City	City
Temporal unit	Quarter	Quarter	Quarter	Quarter

Notes: The methodology underlying the derivation of Table 9 is based on Altonji et al. (2005).

Table 10
Regression results for the placebo treatment estimations.

Column Dependent variable	(1) *SD*	(2) *SD*	(3) *SD*	(4) *P75-P25*	(5) *P75-P25*	(6) *P75-P25*
Starting period of Placebo Treatment	2007Q1	2007Q2	2007Q3	2007Q1	2007Q2	2007Q3
Placebo treatment coefficient	0.001	−0.003	−0.006	−0.002	0.008	0.007
	(0.005)	(0.004)	(0.004)	(0.006)	(0.008)	(0.156)
Control variables	included	included	included	included	included	Included
# of quarters prior and subsequent to the placebo treatment	8 quarters	8 quarters	8 quarters	8 quarters	8 quarters	8 quarters
Spatial unit	City	City	City	City	City	City
Temporal unit	Quarter	Quarter	Quarter	Quarter	Quarter	Quarter

Notes: The coefficient on the placebo treatment variable in the table is estimated under a 19-quarter time-window—including 8 quarters of pre-placebo treatment, 3 quarters of placebo treatment, and 8 quarters of post-placebo treatment. Standard errors are provided in parentheses.

6.4. Placebo and nonparametric tests

In order to further examine our evidence on the effect of information shock on price dispersion, we re-estimate Eq. (2) with a placebo treatment over periods ending prior to the information shock. In particular, we replicate the pre/post treatment structure over a pre-information shock period from 2005Q1 to 2010Q1, specifying time windows of 19 quarters—each of which includes 8 quarters of pre-placebo treatment, 3 quarters of placebo treatment, and 8 quarters of post-placebo treatment—based upon which we re-estimate Eq. (2) by the same method presented in Section 4 above. We use both *SD* and *P75-P25* on the left-hand side of (2). Altogether, we thus estimate Eq. (2) with a placebo treatment 6 times—twice (using *SD* and *P75-P25*) for every window of 19 quarters—during the period 2005Q1–2010Q1, where the variable *Treatment* is equal to 1 for post-placebo treatment periods, and 0 otherwise.

Table 10 presents the estimated coefficient on the *Treatment* variable in the 6 placebo treatment estimations over the period 2005Q1–2010Q1. (Note that the first placebo treatment starts at 2007Q1, following the 8 quarters of pre-treatment period.) It follows that the coefficient on *Treatment* in all of the placebo treatment estimations is insignificantly different from zero. This outcome stands in stark contrast to the significantly negative sign obtained for the coefficient of the real treatment (discussed above). Moreover, note that the coefficient obtained on the placebo *Treatment* variable is about one order of magnitude smaller than the coefficient obtained on the real *Treatment* variable [in absolute value terms—see once again the results from the estimation of Eq. (2) in Table 3].

Finally, we extend the robustness test of our evidence on the effect of the information shock on housing price dispersion to a nonparametric approach. We formally compare the distribution of SD_{tc} (and $P75\text{-}P25_{tc}$) over the 12 (and 4) quarters prior and subsequent to the information shock, applying a Kolmogorov-Smirnov test for equality of distribution functions and a Kruskal-Wallis equality of population test. Outcomes of both tests support the conclusion that price dispersion significantly decreases subsequent to the shock (results are not reported but are available on request).

7. Information shock effect and household and housing market characteristics

In the above analysis, we have provided evidence in support of improved information shock's effect on quality-adjusted price dispersion of housing transactions. We now question whether the effect of the improved information shock varies with household and housing market characteristics.

Specifically, we examine whether the magnitude of the information shock on price dispersion varies across sub-markets that differ in socioeconomic status (SES) and heterogeneity of the housing assets. Intuitively, one may hypothesize either a positive or a negative correlation between the degree of the improved information effect and household SES. On the one hand, one might expect that more privileged households (specifically on the buyer side, as buyers are generally more informationally disadvantaged as compared to sellers; see, e.g., Kurlat and Stroebel, 2015), represented by higher SES, would better exploit the new improved information; and therefore household SES in the market is to be positively associated with the magnitude of the information shock effect on decreased price dispersion. On the other hand, it might be the case that, even when formal price information was unavailable (i.e., prior to the information shock), more privileged households found the means to partially bridge the information gap (by means of, for example, greater access to and improved analysis of asking prices). Hence, when information is formally revealed, higher SES households' marginal benefit from the improved information is relatively limited, as compared to that of less privileged buyers and sellers. The latter—who previously had not only no formal price information available to them, but also experienced very limited access to indirect information channels—now have simple and direct access to all market price information. Hence, their marginal benefit from the price information shock exceeds that of the more privileged households and, accordingly, one can expect a negative correlation between the magnitude of the information shock effect and household SES.

Further, following Kurlat and Stroebel (2015), information friction is of larger concern in areas with greater heterogeneous housing stock. It is thus expected that the price information shock effect on decreased

Table 11
Regression results for the household and housing market characteristic estimation of Eq. (4).

Column	(1)	(2)
Dependent variable	*SD*	*SD*
Constant	0.030**(0.014)	0.025*(0.013)
Treatment (ω_1)	−0.018***(0.004)	−0.018***(0.004)
SES (ω_2)	−0.012**(0.005)	−0.012***(0.005)
Assets (ω_3)	0.341***(0.018)	0.373***(0.017)
SES×*Assets* (ω_4)	−0.028(0.019)	−0.027(0.019)
Treatment ×*SES* (ω_5)	0.009**(0.004)	0.009**(0.004)
Treatment ×*Assets* (ω_6)	−0.036**(0.014)	−0.035**(0.014)
Treatment×*SES*×*Assets* (ω_7)	0.011(0.017)	0.012(0.017)
*CONTROLS*1	Included	Included
$Avg\_Attributes_{ct}$	Included	Included
$SD\_Attributes_{ct}$	Included	Not included
# of observations	4312	4312
# of statistical areas	827	827
R^2 (within statistical areas)	0.26	0.26
Prob>χ^2	0.0000	
# of years prior and subsequent to the treatment	3 years	3 years
Spatial unit	Statistical Area	Statistical Area
Temporal unit	Year	Year

Notes: Table 11 presents results of GLS with random-effects estimation of Eq. (4) with clustered standard errors. Results are robust to the omission of observations in which the *Characteristic* variable is in the top and bottom 5% of its distribution. Standard errors are provided in parentheses. One, two, and three asterisks represent significance at the 10%, 5%, and 1% levels, respectively.

price dispersion is more salient in areas where the housing stock exhibits greater heterogeneity.

Below we report on tests of whether the effect of the improved information shock varies with statistical area measures of household head's SES and housing asset heterogeneity. Based on the statistical area-level sample, we estimate the following [variation of (2) that interacts $Treatment_\tau$ with measures of SES and housing asset heterogeneity]:

$$\begin{aligned} SD_{\tau s} = {}& \omega_0 + \omega_1 Treatment_\tau + \omega_2 SES_s + \omega_3 Assetss_s + \omega_4 SES_s \\ & \times Assets_s + \omega_5 Treatment_\tau \times SES_s \\ & + \omega_6 Treatment_\tau \times Assets_s + \omega_7 Treatment_\tau \times SES_s \times Assets_s \\ & + \vec{\omega}_8 CONTROLS1_{\tau s} + \varepsilon_{6\tau s}, \end{aligned} \tag{4}$$

where SES_s is a statistical area's score on the socioeconomic index, and $Asset_s$ is a measure that summarizes the heterogeneity of the housing stock in a statistical area.[33] Also, $\omega_0 - \omega_7$ and $\vec{\omega}_8$ are estimated parameters and vector of parameters, respectively; $\varepsilon_{6\tau s}$ is a random disturbance term; and all other variables are as described above.[34]

Column 1 in Table 11 presents the outcomes from the estimation of Eq. (4). The results offer evidence of significant variation in the effects of information shock across *SES* and *Assets*. Specifically, the coefficient on *SES* (*Assets*) is negative (positive) and significant at the 5% (1%) level, indicating that house price dispersion decreases with household socioeconomic index (increases with housing asset heterogeneity) in the sub-market. Moreover, the coefficients on the interactions terms of the information shock (*Treatment*) with *SES* and *Assets* are positive and negative, respectively (both significant at the 5% level). Column 2 in Table 11 shows that the outcomes from the estimation of (4) are robust to omitting the control vector of standard deviations of housing asset attributes, $SD\_Attributes_{ct}$, from the right-hand side of (4).

It follows from the outcomes in Table 11 that, prior to the information shock, a higher level of *SES* (*Assets*) is associated with a lesser (greater) price dispersion, *ceteris paribus*. In other words, markets with more privileged households and more heterogenous housing assets exhibit lower and greater price dispersion, respectively, *ceteris paribus*. In addition, considering the interaction terms in (4), results indicate that, while the effect of the information shock decreases overall price dispersion, the magnitude of the effect (rate of change of price dispersion) varies by the socioeconomic index and asset heterogeneity in the market. Specifically, for relatively high (low) levels of *SES*, increasing the level of *Assets* is associated with a greater (smaller) information shock effect on the rate of decreased price dispersion [that is, the effect of the shock on decreased price dispersion increases (decreases) with asset heterogeneity in markets with more (less) privileged households]; while for any given level of *Assets*, increasing the level of *SES* is associated with a smaller information shock effect on the decreased price dispersion (that is, markets with less privileged households exhibit a greater information shock effect on decreased price dispersion for any level of housing asset heterogeneity).

The implications of our findings are twofold. First, the effect of improved information shock on price dispersion is most pronounced in markets where households exhibit relatively low levels of socioeconomic characteristics. This is consistent with the notion that market participants in less privileged areas, having limited access to means that may ameliorate information shortage, are experiencing a greater effect of the improved public information, while in areas with a higher socioeconomic population, transactions involve greater information even when the latter is not formally available. Furthermore, results indicate that while all sub-markets experience the improved information effect, the magnitude of the effect is positively correlated with the extent of information frictions—as characterized by the degree of housing asset heterogeneity—specifically, in areas with more privileged households.

[33] We derive $Asset_s$ by re-estimating Eq. (1) for all s—a total of 829 estimated equations. However, for each s, we then compute the projected log price based only on asset characteristics: $\widehat{\ln(P_{i\tau s})} = \hat{\theta}_{0,s} + \vec{\hat{\theta}}_{1,s} CHARACTERISICS_{i\tau s}$. That is, $\widehat{\ln(P_{i\tau s})}$ is the projected log price that follows from the estimation of (1) net of $\vec{\hat{\theta}}_{2,s} TFE_\tau + \hat{\varepsilon}_{5i\tau s}$. Given this step, the term $\widehat{\ln(P_{i\tau s})}$ summarizes only the difference in quality of all assets transacted in s over the examined period. Finally, computing the standard deviation of $\widehat{\ln(P_{i\tau s})}$ across all i in s generates $Asset_s$, a measure of housing asset heterogeneity in s.

[34] As SES_s and $Assets_s$ are specific to a statistical area s, we omit the vector of statistical area fixed-effects (LFE_s) from the estimation of (4). Also, the outcomes on the association between the information shock effect and *SES* are robust to replacing households' *SES* measure with household heads' measures of either education or income in the statistical area. These results are not reported but are available upon request.

8. Summary and conclusions

This research provides new empirical evidence on the effect of information shock on quality-adjusted housing price dispersion. The analysis examines a unique Israeli experience wherein the Tax Authority was court-ordered to publicly disclose information on all past and present real estate transactions.

Statistical findings provide evidence in support of improved information effect on the dispersion of transaction prices. We find that standard deviation of quality-adjusted prices has decreased by about 18% subsequent to the improved information shock. Further, we find evidence that the magnitude of the information effect varies with individual (socioeconomic status) and market (asset heterogeneity) characteristics. Outcomes are robust to a series of sampling and test-design issues. Research findings provide real-world evidence suggesting the substantial effects of price transparency in a market where transactions involve significant and long-term individual economic consequences.

Appendix

Derivation of control variables $SD\_\hat{P}_{tc}$, $\Delta\hat{P}_{tc}$, $\Delta\hat{P}_{\tau s}$, and $SD\_\hat{P}_{\tau s}$:

For each city c in the sample we estimate:

$$\ln(P_{itc}) = \beta_{0,c} + \beta_{1,c}CHARACTERISTICS_{itc} + \beta_{2,c}TFE_t + \varepsilon_{2itc} \text{ for all } c, \tag{A1}$$

where the only difference between Eqs. (A1) and (1) is that *TFE* in (A1) is a vector of monthly (rather than quarterly) time fixed-effects. All other variables are as described above.

Eq. (A1) is estimated for each city c (altogether 57 estimations whose average R^2 is 0.80, with a minimum of 0.63 and a maximum of 0.90). We use the *TFE* coefficients to produce a price index for each city by which we compute the 6-month (ending at t) rate of change in quality-adjusted housing prices in city c, $\Delta\hat{P}_{tc}$, and the standard deviation (over a 6-month period ending at t) of the rate of change in the quality-adjusted price, $SD\_\hat{P}_{tc}$—both appear on the right-hand side of Eq. (2). Similarly, we use the derived price index for each city to compute the 12-month rate of change in quality-adjusted housing prices in city c for each year, $\Delta\hat{P}_{\tau s}$, and the standard deviation of the rate of change in the quality-adjusted price in city c for each year, $SD\_\hat{P}_{\tau s}$—both appear in the statistical- (rather than city-) level estimations.

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