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THE ROLE OF AUCTIONS AND NEGOTIATION IN HOUSING PRICES

Abstract

Using Sydney and Melbourne transactions, we show that how properties sell matters for housing price dynamics. Auction prices forecast better and display much less momentum than negotiated prices. This is consistent with the two mechanisms transmitting buyer vs. seller shocks to prices differently and, in light of auction and bargaining theories, suggests the source of momentum is sluggishness in sellers' valuations. Other explanations, such as differences in precision, slow diffusion of shocks among buyers, or endogenous selection of the sales mechanism, fail to explain our findings. Our estimates also indicate that sellers have at most equal bargaining power in negotiations.

JEL Classification: L11, D44, R31

Keywords: auctions, Bargaining, housing prices

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The Role of Auctions and Negotiation in Housing Prices

By David Genesove and James Hansen*

Draft: July 12, 2016

Using Sydney and Melbourne transactions, we show that how properties sell matters for housing price dynamics. Auction prices forecast better and display much less momentum than negotiated prices. This is consistent with the two mechanisms transmitting buyer vs. seller shocks to prices differently and, in light of auction and bargaining theories, suggests the source of momentum is sluggishness in sellers' valuations. Other explanations, such as differences in precision, slow diffusion of shocks among buyers, or endogenous selection of the sales mechanism, fail to explain our findings. Our estimates also indicate that sellers have at most equal bargaining power in negotiations.

JEL: D44, D49, R30, R32

Keywords: Bargaining, Auctions, Real-estate pricing

The last financial crisis made apparent the importance of housing market dynamics. However, these dynamics are not easily reconciled to the usual models. Perhaps most resistant to explanation is the highly positive autocorrelation in price growth (momentum). First observed by Case and Shiller (1989) for US single family homes and a repeated finding across countries and time,[1] this phenomenon is at odds with a standard asset model for housing markets. As noted by Glaeser, Gyourko, Morales and Nathanson (2014), "The model fails utterly at explaining the strong, high frequency positive serial correlation of price changes."

A number of recent papers have attempted to model housing price dynamics, incorporating search frictions (Capozza, Hendershott and Mack (2004), Caplin and Leahy (2011), Díaz and Jerez (2013) and Head, Lloyd-Ellis and Sun (2014)),

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[1] See Titman, Wang and Yang (2014) for a more recent study showing this empirical regularity.

adaptive expectations (Sommervoll, Borgersen and Wennemo (2010)), momentum traders (Piazzesi and Schneider (2009)), and kinked demand curves (Guren (2015)). Yet, these papers have had limited success in generating the high degree of positive autocorrelation. Head, Lloyd-Ellis and Sun (2016), for example, explains less than half of the first autocorrelation coefficient in price growth and none of the second, while Díaz and Jerez's (2013) model generates no autocorrelation at all.

The first contribution of this paper is to show that price growth autocorrelation is much smaller or even absent for homes sold via auctions rather than the usual bilateral negotiations. To do so, we use Australian data, where auctions are a substantial fraction of housing transactions. Using 1992 to 2012 sales data for Sydney and Melbourne (around 40 per cent of all Australian sales), we find that the autocorrelation properties of prices determined through bilateral negotiation (hereafter private-treaty prices) and auction (hereafter auction prices) are very different. Although private-treaty price growth is highly autocorrelated, as previous studies find, auction price growth is much less so. Indeed, we cannot reject the null that auction prices are informationally efficient and follow a random walk with drift. We then exploit the differential structure of auctions and negotiations to argue that the momentum that we observe in private-treaty prices, and by extension the momentum that has been observed by previous researchers, reflects a sluggish seller response. Sellers who are slow to respond to market conditions play a part in generating autocorrelation in both Caplin and Leahy's (2011) and Guren's (2015) models, but neither paper presents evidence in support of the assumption.

In addition to being informationally efficient, we find auction prices useful for predicting future housing prices. This is true for both private-treaty prices and average prices overall. It is consistent with auction prices quickly updating in response to changes in a common stochastic trend in housing prices. In contrast, we find that private-treaty prices are useful for predicting neither auction prices nor average housing prices overall. Notably, private-treaty prices only fully reflect changes in the common stochastic trend with a lag of almost a year. Moreover, private-treaty prices are sensitive to temporary shocks specific to them. These findings are striking given that auctions make up less than 17 per cent of all transactions on average so that, a priori, one would expect private-treaty prices to be the more informative measure.

Why such large differences in efficiency and information content? We argue that these findings are consistent with asymmetry in the weighting of buyers' and sellers' valuations across the two sales mechanisms, and a slow adjustment by sellers to market conditions.

Private-treaty sales are typically the result of bilateral negotiation between one buyer and one seller.[2] In standard complete or incomplete information bargaining models, both seller and buyer values influence price.[3] Indeed, with equal bargaining weights, which our results support, and buyers and sellers drawing independently from equally dispersed uniform distributions, shifts in the support of either distribution have the same effect on the expected price.

Auctions are different. In an auction, such as the open-outcry (English) auction used in Sydney and Melbourne, a seller offers the property to many buyers, who simultaneously compete through a bidding process. In the absence of a seller reserve, auctions are solely determined by the distribution of buyers' valuations. Even when a seller reserve is used and so the auction may not always be successful, shocks to sellers' valuations have a smaller effect on price than buyers' valuations. This is true for the uniform distributions case that we present in Section IV.A and for a wide range of distribution pairs that we consider in the Appendix, calibrated to match the sale rates (hereafter clearance rates) observed in these cities.

The scope for slower seller adjustment to changing market conditions is supported by a number of housing market phenomena. These include: the greater cyclicality of sales than housing prices (Leamer (2007)); the lower seller time on the market in 'hot' markets (a stylized fact for Wheaton (1990) and Krainer (2001), and documented in Genesove and Han (2012))[4]; the positive correlation between the transaction to list price ratio and short run demand growth (Genesove and Han (2012)), and the positive correlation between that ratio and unexpected or expected price growth (Haurin et al. (2013)). We provide additional supporting evidence from our data: list prices, a direct measure of sellers reservation values,

[2] If there are bidding wars, other buyers' values will also affect private-treaty prices. This can be interpreted as misclassification of some private-treaty sales that would be better thought of as auctions. It would imply that we are underestimating the extent of the difference in the behaviour of (true) private-treaty and auction prices.

[3] For complete information, see Nash (1950), Harsanyi (1956), Cross (1969) and Roth (2012), and Fudenberg and Tirole (1991) and Napel (2002) for reviews of the literature. For incomplete information, see Ausubel, Cramton and Deneckere (2002) and references cited therein, as well as Čopič and Ponsatí (2008).

[4] Head et al. (2014), however, provide a model in which seller time on the market anti-cyclicality arises instead out of the short run fixity of housing and demand shock persistence.

lag both auction and private-treaty prices, and a 'Phillips-curve' governs the relationship between price growth and the clearance rate. List prices also allow us to estimate seller bargaining power: a relatively precise 0.5 for Sydney, while one insignificantly different from both equal and no bargaining power for Melbourne.

Why sellers' values lag buyers' is beyond the scope of our investigation, but we make some brief comments. First, the asymmetric matching institution, in which sellers list homes and prices, while buyers do not list their preferences or even their identities, makes seller search more public than buyer search. Consequently, information on shocks to sellers may diffuse quickly through their listing, de-listing and list price decisions, becoming common information to sellers and buyers alike, while buyer shocks only become publicly known when actualised in transacted prices and those prices publicised. On the other hand, information flows during the listing period may more easily be absorbed by buyers. Whereas sellers may choose to be passive once they have listed their property, allowing an agent to represent them in dealings with visiting buyers, buyers are almost always active in visiting properties themselves.[5]

Second, the higher dimensionality of the buyer problem, which includes not only price but also the home attributes, which are fixed on the seller side, forces the buyer to constantly reassess willingness to pay cross-sectionally. With psychological or information gathering or decision costs already incurred, buyers may be more prepared than sellers to reassess their valuations over time as the market changes. Third, buyers moving into an area may be more attuned to changes in future housing services values than sellers, who, on net, are moving out (Leamer (2007)). As Guren (2015) has noted, if buyer arrivals are concave in the seller's list price, strategic complementarity assures that these various phenomena need not necessarily affect a large fraction of sellers to have large effects on sellers' list prices, and through them, transaction prices. Fourth, equity lock-in and loss aversion, which explain seller price rigidity in downturns, appear less relevant here because prices in our data are generally increasing.[6]

We also consider other explanations for the joint behaviour of prices. One is that by incorporating information from more than one buyer, auction prices more

[5] Differential information flows or asymmetry between buyer and seller behaviour has been emphasised in previous research, see for example Anenberg (2011) and Berkovec and Goodman (1996).

[6] See Stein (1995), Genesove and Mayer (1997, 2001), Engelhardt (2003), and Anenberg (2011).

precisely estimate an underlying common-value component in buyers' valuations.[7] Common values arise endogenously in search environments with uncertainty over market conditions, as Merzyn, Virag and Lauermann (2010) stress. Being forward looking, the value of continued buyer search should be a good predictor of future prices. In addition, auction theory suggests that the winning bid at an auction will reflect the common-value component given a sufficient number of bidders – converging to it if that is the only component of buyers' valuations and to a function of it if there is a private-value component as well.

Notwithstanding the theoretical appeal of this explanation, price indices are, nonetheless, formed by averaging prices across many transactions. Although an individual auction may provide a more accurate measure of the underlying state variable, the average auction price may not perform better than the average private-treaty price. There are seven (Melbourne) to ten (Sydney) times as many private-treaty transactions as auctions. Thus, a lesser precision in a private-treaty price from incorporating fewer signals of the common value could be offset by the larger set of signals incorporated into the average price through more transactions. We find that the number of transactions is so large relative to price dispersion at the individual transaction level, as measured by the root mean squared error of the underlying hedonic regressions, that aggregation effectively offsets any precision gains that might originate at the transaction level.

Yet another difference between auctions and negotiations is that the former draws its price from the right tail of the buyer distribution. A diffusion over time of common buyer shocks through the buyer distribution will lead to a lead-lag relationship between auction and private-treaty prices. We argue, however, that the predicted relationship is different from what we observe.

Another possible explanation is that auction transactions garner greater publicity than negotiated transactions, being more dramatic, attended by more people, and having their results published in newspapers and auction company websites. If market participants form expectations conditioning on observed past transactions, the publicity given to those previous transactions will matter.[8] We find support for

[7] See Kremer (2002), for example, which establishes this result using limiting arguments.

[8] We have the full set of transactions, and date them according to the date of transaction and not publication. A related issue is the distinction between the contract date and the settlement date. However, the difference between the two is very similar on average for both sale mechanisms and in both cities.

this explanation only if we assume that sellers alone use lagged auctions information – otherwise it implies that auction and private-treaty prices have more similar autocorrelation properties than they do.

Finally, we consider whether our results are sensitive to the measurement of prices, endogenous selection of the sale mechanism, and the characteristics and location of homes sold. Using alternative measures of price, including fewer attribute controls to maximise sample size, or using repeat-sales indices to better control for unobserved attributes, has little effect on our findings.[9] Nor does the use of price indices that adjust for endogenous selection. Focusing on within-group variation, first within sub-city districts and then by the type of homes sold, has little effect either. Even at the district level or by home type, auction prices remain locally informative and Granger cause private-treaty prices, but the reverse is not true.

Australia stands out for its non-trivial share of non-foreclosure auctions, and so is particularly useful for investigating price formation. Our findings should also be of interest for other countries because of the increasing frequency of auctions or bidding wars in housing markets elsewhere (Han and Strange (2014)). Note that we are not concerned with the relative profitability of auctions versus negotiations – of interest in a number of theoretical papers (Bulow and Klemperer (1996, 2009)) – unless this changes in a predictable fashion correlated with price dynamics.

Interest in housing price formation and the forecastability of prices stems from both macro and micro policy concerns. At the macro level, mortgage performance, the solvency and stability of the banking system, household collateral levels and investment and saving, all depend on changes in housing prices (Iacoviello (2005), Iacoviello and Neri (2010)).[10] The dramatic run-up in prices and subsequent falls in many countries played a key role in the global financial crisis, and has generated wider interest in housing prices dynamics.

This paper should also help in understanding the link between sales mechanisms and price formation. A large theoretical literature compares outcomes such as efficiency, seller revenue and information aggregation across mechanisms, especially

[9] Using fewer attributes increases the autocorrelation of price growth, consistent with serially correlated changes in the composition of homes sold (Hansen (2009)). However, it has little effect on the relative information content of auction vs. private-treaty prices and auction price growth remains much less autocorrelated than private-treaty price growth.

[10] Case, Shiller and Quigley (2005) argue that there are large wealth effects operating between the value of homes and consumption.

auctions (see for example, Bulow and Klemperer (1996, 2009), Kremer (2002)), but also between them and posted prices (Wang (1993, 1998)). An empirical literature compares price levels across different mechanisms, especially on the Internet (e.g., Lucking-Reiley (1999), Einav et al. (2015)). Most theoretical and the empirical work has a single transaction focus. This paper contributes by providing empirical evidence on how different selling mechanisms map changes in the underlying distribution of buyers' and sellers' valuations into average price changes over time.

The next section discusses the data and construction of the price indices. Sections II and III discuss the differences in autocorrelation between the two price measures, their relative information content when forecasting future price growth, and their sensitivity to permanent and temporary shocks. Section IV interprets our findings in the light of alternative theories of price formation and the final section concludes.

I. Data and Measurement

Our primary data source is a census of all housing sales in Sydney and Melbourne between 1992:I and 2012:IV, about 40 per cent of all Australian housing sales over that period. Provided by Australian Property Monitors (APM), this is an update of data previously used by Prasad and Richards (2008) and Hansen (2009).[11]

Private-treaty is the most common mechanism used for selling housing in these two cities. Sales where an auction mechanism was used (or planned to be used) as part of a successful sale make up around 12 per cent of the Sydney sample and 17 per cent for Melbourne (Table 1, columns one and two).

Table 1—Overview of Sales Mechanisms Used

	Percentage of total observations[(a)]		**Percentage filtered for analysis[(b)]**	
Transaction type	Sydney	Melbourne	Sydney	Melbourne
Pre- or post-auction	2.73	3.72	na	na
Sold at auction	8.83	13.01	9.30	13.90
Private treaty	88.46	83.26	90.70	86.10
Auction frequency	11.56	16.73	9.30	13.90
Total observations	1 763 032	1 677 925	1 652 585	1 498 549

Note: [(a)]Percentage of total observations where an auction was used (or planned to be used) as part of a successful sale; [(b)]percentage of observations after removing identified pre- and post-auction sales, private-treaty sales where an auction was used in the 90 days prior to the exchange of contracts.

[11]In providing these data, APM relies on a number of external sources. These include the NSW Department of Finance and Services for property sales data in Sydney and the State of Victoria for property sales data in Melbourne. For more information about these data, see the Copyright and Disclaimer Notices in the Appendix.

In the following, we restrict attention to properties sold at auction when measuring auction prices (Table 1, columns three and four) and properties sold via bilateral negotiation, with no prior offering (within the previous 90 days) at auction, when measuring private-treaty sales.[12] Using hedonic price regressions similar to those discussed below, the average conditional price difference between a property sold through an auction and through a private-treaty is 4.2 per cent for Sydney and 5.1 per cent for Melbourne.[13]

To compute the indices we use hedonic price regressions. At the city-wide level, Hansen (2009) has shown that hedonic regressions provide an accurate estimate of the composition-adjusted price change in housing. The specification is

$$\ln P_{ijt} = \sum_{t=0}^{T} \gamma_t D_{it} + \sum_{j=1}^{J} \beta_j PC_{ij} + \sum_{k=1}^{K} \theta_k C_{ikt} + \varepsilon_{ijt}$$

The variable $\ln P_{ijt}$ is the logarithm of the sale price for property i, in postcode j and at time t; D_{it} is a time dummy equal to 1 if sold in quarter t and zero otherwise; PC_{ij} is a postcode dummy ; and C_{ikt} is the kth attribute of the home at time t. For each city, we run separate regressions for auctions and private-treaty sales.

The attributes are the number of bedrooms, the number of bathrooms, the logarithm of a measure of the size of the property, the type of the home sold (house, semi-detached, terrace, townhouse, cottage, villa, unit, apartment, duplex, studio) and the interaction of the property type with each of the first three variables.[14] When estimating recursively to generate out-of-sample forecasts, we use the maximal sample size and include the property-type control only. When generating in-sample estimates, we include all controls and their interaction effects for Sydney, but only include the property-type control for Melbourne unless stated otherwise. Our estimates span 1992:I (1993:I) to 2012:V for Sydney (Melbourne).

Figure 1 reports, for each city, two-quarter-ended annualised growth of separate hedonic price indices for auction, private-treaty and all-sales prices. Although highly correlated, the three indices are not fully synchronised, with auction prices leading all-sales and private-treaty prices, most notably around turning points.

[12] See Table 1, Note (b).

[13] This is measured using an additional dummy variable for whether the property is sold via auction or private-treaty.

[14] For houses, size is the total land area in square metres. For units or apartments, it is typically a measure of the building area, but can also be the internal area depending on the data source.

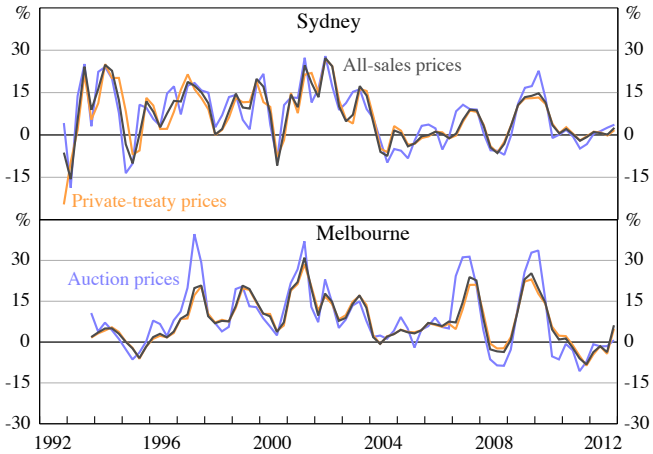

FIGURE 1. COMPARISON OF AUCTION, PRIVATE-TREATY AND ALL-SALES PRICES: TWO-QUARTER-ENDED ANNUALISED GROWTH

II. Prediction

In this section we examine three questions:

1) Do auction prices and private-treaty prices have different autocorrelation properties?

2) Do they perform differently when predicting out-of-sample?

3) Do they perform differently when predicting one another in-sample?

The first question speaks to the well-established literature on the efficiency of housing markets, which shows that housing price growth is positively autocorrelated (Case and Shiller (1989), Cutler, Poterba and Summers (1991), Cho (1996) and Capozza, Hendershott and Mack (2004) among others). Differences in momentum allow us to discriminate between alternative models of housing market dynamics. The second addresses whether gains in predictive content are available in real time.

We also consider in-sample analysis for three reasons: focusing on the full sample of data avoids the effects of revisions to the estimated price indices that could affect out-of-sample forecasting; it allows us to relax the finite lag VECM representation assumption maintained in the out-of-sample analysis;[15] and out-of-sample analysis can entail a loss of information and power (Inoue and Kilian (2005)).

[15] Although a VECM with finite lags is a natural framework for modelling prices given that they are likely to share the same common trend, it is not an immediate implication of theory. In Section IV we build up a case to support this representation, rather than assume it is valid.

A. Momentum

The autocorrelation functions in Figure 2 show that all-sales price growth is positively autocorrelated for up to one year, but that the strongest correlations are for the first two quarterly lags. All of the positive autocorrelation in aggregate price growth for Sydney arises from private-treaty prices; no evidence suggests that auction price growth is positively autocorrelated. Indeed, auction prices follow a random walk with drift. This striking result suggests that all available information concerning housing prices is fully incorporated into auction prices within a quarter.

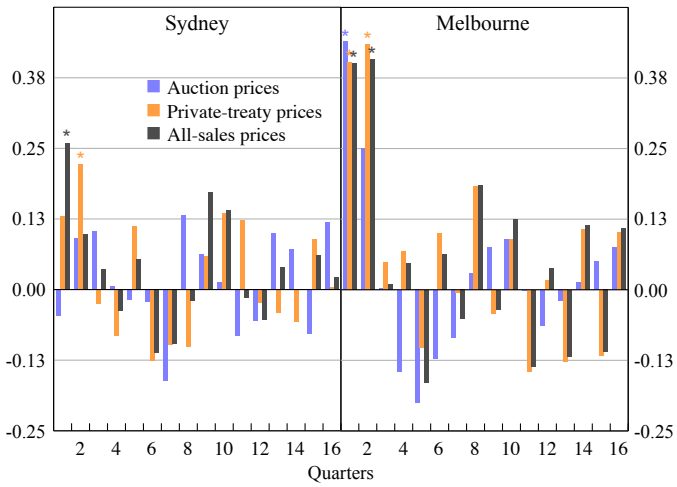

FIGURE 2. AUTOCORRELATION FUNCTIONS FOR PRICES GROWTH

Note: Asterisks denote significance at the 5 per cent level when using Bartlett's MA(q) formula.

For Melbourne, most of the autocorrelation in all-sales price growth is also driven by private-treaty price growth, although there is some weak evidence of first-order autocorrelation in auction price growth.[16] The difference in autocorrelation functions, with private-treaty price growth being more autocorrelated than auction price growth, will subsequently be useful for determining whether buyer or seller valuations are sluggish, as discussed further in Section IV.

B. Out-of-sample

We now consider whether price indices conditioned on the sale mechanism are useful for predicting all-sales price growth in real time. Specifically, we consider whether including lagged auction prices or lagged private-treaty prices improves

[16] The autocorrelation function for Melbourne auction price growth is even smaller if one includes detailed attributes data when estimating the hedonic price indices and using the sample from 1997:IV onwards.

upon the one-quarter-ahead forecasts of all-sales price growth in a single equation autoregressive model. We compare the following three forecasting models[17]

$$(1) \qquad \Delta s_t = \mu_s + \sum_{j=1}^{J} \phi_j \Delta s_{t-j} + \varepsilon_t^s$$

$$(2) \qquad \Delta s_t = \mu_s + \Gamma_s s_{t-1} + \Gamma_a a_{t-1} + \sum_{j=1}^{J} \phi_j \Delta s_{t-j} + \sum_{j=1}^{J} \gamma_j^a \Delta a_{t-j} + \varepsilon_t^{s,a}$$

$$(3) \qquad \Delta s_t = \mu_s + \Gamma_s s_{t-1} + \Gamma_p p_{t-1} + \sum_{j=1}^{J} \phi_j \Delta s_{t-j} + \sum_{j=1}^{J} \gamma_j^p \Delta p_{t-j} + \varepsilon_t^{s,p}$$

where s_t is the all-sales housing price index, a_t is the auction price index and p_t is the private-treaty price index. Equation (1) is the benchmark model, a univariate autoregression in average all-sales price growth. Equation (2) adds lags in auction prices, and allows all-sales and auction prices to be cointegrated, consistent with the price indices sharing a common stochastic trend. Equation (3) incorporates lags of private-treaty prices instead of lags of auction prices and also allows for cointegration. Evidence for cointegration is provided in the Appendix (Section VII, Table 24), though similar results are obtained below without this assumption.

To measure out-of-sample prediction accuracy, we define

$$\sigma_i^2 \equiv E\left(\hat{s}_{t+1|t}^i - s_{t|t} - \left(s_{t+1|t+1} - s_{t|t+1}\right)\right)^2$$

for $i = 1, 2, 3$ as the respective mean-squared prediction errors (MSPEs) for one-quarter-ahead all-sales price growth associated with Equations (1), (2) and (3) respectively. $\hat{s}_{t+1|t}^i \equiv E\left(s_{t+1}^i \mid I_t\right)$ is the one-quarter-ahead forecast of the log all-sales price level based on Equation i (for $i = 1, 2, 3$) and using the information available at time t. $s_{t|\tau}$ is the measured value of the log all-sales price level at time t given all available information up to time $\tau \geq t$. We consider whether the MSPEs are statistically different among (1), (2) and (3) using pairwise comparisons and the MSE-t test statistic discussed in McCracken (2007).[18]

[17] In these, and all subsequent out-of-sample forecasting tests, we use four lags when using Sydney data and three lags when using Melbourne data. This is based on likelihood-ratio and residual serial correlation tests, as well as information criteria. For Melbourne, quarterly seasonal dummies are included as additional control variables, consistent with evidence of seasonality in Melbourne.

[18] This test statistic is equivalent to the S_1 test statistic proposed by Diebold and Mariano (1995). We use the critical values tabulated in McCracken (2007), which notes that S_1's distribution may be ill approximated by a normal distribution for nested prediction equations. Similar results are obtained using

Table 2 shows that Equation (2) outperforms the benchmark model: there is information content in lagged auction prices. In both cities, the MSPEs for (2) are significantly lower relative to the benchmark model by about 10 and 18 per cent for Sydney and Melbourne (column one, rows one and three). In contrast, little suggests that private-treaty prices improve upon the benchmark model forecasts; the null that the forecast accuracy of Equation (3) is the same cannot be rejected at conventional significance levels (column one, rows two and four).

Table 2—Pairwise Nested Model MSPE Comparison

	$\frac{\sigma^2_{y\in\{a,p\}}}{\sigma^2_s}$	MSE-t statistic
Sydney		
$H_0 : \sigma^2_s - \sigma^2_a = 0$	0.90**	0.85
$H_0 : \sigma^2_s - \sigma^2_p = 0$	0.93	0.26
Melbourne		
$H_0 : \sigma^2_s - \sigma^2_a = 0$	0.82**	1.46
$H_0 : \sigma^2_s - \sigma^2_p = 0$	0.97	0.17

Note: The alternative hypothesis for each test is that the MSPE of the restricted model, σ^2_s, is greater than the unrestricted alternative (either σ^2_a or σ^2_p); recursive estimation is used starting with the sample period from 1992:I to 2007:I for Sydney and from 1993:I to 2008:III for Melbourne; ***, ** and * denote significance at the 1, 5 and 10 per cent levels.

To further investigate which of auction prices or private-treaty prices contains more information about future price growth, we consider whether these price indices are useful in predicting one another using out-of-sample Granger causality tests. Again, allowing for auction and private-treaty price to share a common stochastic trend, the unrestricted model used for our tests is given by

$$\Delta a_t = \mu_a + \alpha_a (a_{t-1} - \beta p_{t-1}) + \sum_{j=1}^{J} \Gamma_j^{aa} \Delta a_{t-j} + \sum_{j=1}^{J} \Gamma_j^{ap} \Delta p_{t-j} + \varepsilon_t^a \tag{4}$$

$$\Delta p_t = \mu_p + \alpha_p (a_{t-1} - \beta p_{t-1}) + \sum_{j=1}^{J} \Gamma_j^{pa} \Delta a_{t-j} + \sum_{j=1}^{J} \Gamma_j^{pp} \Delta p_{t-j} + \varepsilon_t^p \tag{5}$$

The null hypotheses are that auction prices do not Granger cause private-treaty prices, $H_0 : \alpha_p = \Gamma_j^{pa} = 0$ for all j, and that private-treaty prices do not Granger cause auction prices, $H_0 : \alpha_a = \Gamma_j^{ap} = 0$ for all j. The McCracken (2007) and Clark and West (2007) tests of these hypotheses in Table 3 reject the null that auction prices do not Granger cause private-treaty prices in both cities, but fail to reject

Clark and West's (2007) MSPE-adj t statistic.

the null that private-treaty prices do not Granger cause auction prices in Sydney (and only find weak evidence to reject it in Melbourne). These results confirm that auction prices are more useful, when forecasting out-of-sample.

Table 3—Out-of-sample Granger Causality Tests

	Sydney	Melbourne
H_0 : Auction prices do not Granger cause private-treaty prices		
MSE-t	1.55***	1.38**
MSPE-adj t	2.82***	2.34***
H_0 : Private-treaty prices do not Granger cause auction prices		
MSE-t	–1.06	0.63*
MSPE-adj t	0.50	1.46*

Note: ***, ** and * denote significance at the 1, 5 and 10 per cent levels of significance respectively; MSE-t is the Diebold and Mariano test statistic used in the context of a nested model forecast comparison as discussed in McCracken (2007); MSPE-adj t is an alternative test statistic proposed by Clark and West (2007); estimates and out-of-sample forecasts are generated recursively with the initial in-sample estimation period from 1992:I to 2002:III for Sydney, and from 1993:I to 2002:III for Melbourne.

C. *In-sample*

Although out-of-sample findings are informative for assessing real time forecasting performance, they can entail information and power losses relative to in-sample comparisons (Inoue and Kilian (2005)), and can be affected by revisions. Accordingly, Table 4 revisits the causality tests using the in-sample approach of Toda and Yamamoto (1995) and includes all attributes data.[19] The first four rows support the previous findings. For both cities, they reject the null that auction prices do not Granger cause private-treaty prices, but are unable to reject the null that private-treaty prices do not Granger cause auction prices.

We conduct two further in-sample specification checks. The first shows that auction prices follow a random walk with drift and cannot be explained using lagged price information (Table 4, rows (5)–(6)).[20] That is, auction prices are informationally efficient. This is a striking finding given previous evidence of the highly positive autocorrelation found in price growth across countries and time. For private-treaties, there is clear evidence of informational inefficiency – lagged auction and private-treaty prices are useful in predicting them (Table 4, rows (7)–(8)).

The second, shown in Table 5, checks whether the error-correction specification of the auction–private-treaty price relationship, makes economic sense. Consistent

[19] Conditioning on the assumption of cointegration provides similar results.

[20] This result is also confirmed using a univariate test that regresses auction price growth on lags of auction price growth (i.e. imposing the restrictions that private-treaty prices do not Granger cause auction prices and that auction prices are I(1) in levels). The p-value for Sydney (Melbourne) is 0.35 (0.13).

Table 4—In-sample Granger Causality Tests

	All controls	
Null hypothesis	Sydney	Melbourne
$a_t \overset{GC}{\nrightarrow} p_t$	69.96***	12.57***
	(0.00)	(0.01)
$p_t \overset{GC}{\nrightarrow} a_t$	5.59	3.70
	(0.35)	(0.45)
$E\left(a_t \mid \mathcal{I}_{t-1}^{a,p}\right) = a_{t-1}$	7.60	11.84
	(0.57)	(0.11)
$E\left(p_t \mid \mathcal{I}_{t-1}^{a,p}\right) = p_{t-1}$	79.47***	31.99***
	(0.00)	(0.00)

Note: ***, ** and * denote significance at the 1, 5 and 10 per cent levels. $\overset{GC}{\nrightarrow}$ is a test for non-Granger causality; a_t is the auction price, p_t the private-treaty price and $\mathcal{I}_{t-1}^{a,p}$ is the information set at time $t-1$ (conditioning on lagged auction and private-treaty prices). All controls includes the property type and interactions with the number of bedrooms, the number of bathrooms, and the logarithm of the size of the property. For Melbourne this sample is restricted to 1997:IV onwards and includes controls for seasonality; p-values are in parentheses.

with our previous findings, while private-treaty prices respond positively to the lagged deviation between auction and private-treaty prices, auction prices do not respond to it. Normalising on auction prices, the cointegration parameters, β, also look reasonable and not too far from 1, as expected.

Table 5—Cointegration and Adjustment Parameter Estimates

	Auction prices		Private-treaty prices	
Sydney				
Cointegration parameter		1	$-\beta$	-1.05***
		(.)		(0.01)
Adjustment parameter	α_a	–0.10	α_p	0.42***
		(0.23)		(0.15)
Melbourne				
Cointegration parameter		1	$-\beta$	-1.08***
		(.)		(0.01)
Adjustment parameter	α_a	–0.02	α_p	0.18**
		(0.14)		(0.07)

Note: Cointegration and adjustment parameter estimates are obtained using Johansen MLE and normalising the coefficient on auction prices to 1; ***, ** and * denote significance at the 1, 5 and 10 per cent levels respectively and are with respect to 0 for the adjustment parameters and 1 for the cointegration parameters corresponding to Equations (4) and (5); standard errors are reported in parentheses.

In sum, the data are consistent with the following facts:

1) Forecasts from an autoregression of all-sales prices can be significantly improved upon by including lagged auction price information. Lagged private-treaty prices are less informative.

2) Auction prices Granger cause private-treaty prices, but the reverse is not true. This holds both in- and out-of-sample.

3) Auction prices are informationally efficient. Private-treaty prices are not.

4) Auction and private-treaty prices are cointegrated.

III. The Persistence of Shocks

We now examine the persistence of shocks to housing prices exploiting the differential behaviour of price by the mechanism of sale. We use two conditions supported by the data: (a) auction and private-treaty prices are cointegrated and can be represented by a VECM; and (b) private-treaty prices do not Granger cause auction prices. Together, these imply the following model for auction and private-treaty price growth

$$\Delta a_t = \mu_a + \sum_{j=1}^{J} \Gamma_j^{aa} \Delta a_{t-j} + \varepsilon_t^a \tag{6}$$

$$\Delta p_t = \mu_p + \alpha_p \left(a_{t-1} - \beta p_{t-1}\right) + \sum_{j=1}^{J} \Gamma_j^{pa} \Delta a_{t-j} + \sum_{j=1}^{J} \Gamma_j^{pp} \Delta p_{t-j} + \varepsilon_t^p \tag{7}$$

These conditions are also sufficient for identifying the effects of permanent and temporary shocks to auction and private-treaty prices.[21] A permanent shock is defined as having an effect on long-run forecasts of auction and private-treaty prices whereas a temporary shock has no such effects.

Table 6 reports forecast error variance decompositions of the permanent and temporary shocks, assuming both conditions hold. We only report them for private-treaty prices, as the conditions jointly imply that shocks to auction prices are all permanent. Relaxing (b) by assuming only that the long-run adjustment parameter in the auction price equation is zero yields almost identical results.[22]

Almost half of the forecast error variation in private-treaty prices one-quarter-ahead is due to temporary shocks (columns two and four). At the two (four) quarter-ahead horizon, they account for about 20 (10) per cent of the variance.

Figure 3 graphs estimates of the permanent and temporary (transitory) shocks

[21] See Fisher and Huh (2007) and Pagan and Pesaran (2008).
[22] That is, only imposing $\alpha_a = 0$ rather than $\alpha_a = \Gamma_j^{ap} = 0$ for all $j = 1, ..., J$ in Equation (4).

Table 6—Forecast Error Variance Decompositions for Private-treaty Prices

Forecast horizon	**Sydney**		**Melbourne**	
	Permanent	Temporary	Permanent	Temporary
1	0.45	0.55	0.53	0.47
2	0.77	0.23	0.81	0.19
3	0.82	0.18	0.87	0.13
4	0.89	0.11	0.91	0.09
32	0.99	0.01	0.99	0.01

over time. The former are clearly much larger than the latter. The second half of the 1990s and early 2000s stands out for substantial permanent shocks, consistent with the typical explanations for changes in housing prices during this period, including financial deregulation, productivity improvements, long-run shifts towards greater credit availability, and lower real interest rates.[23]

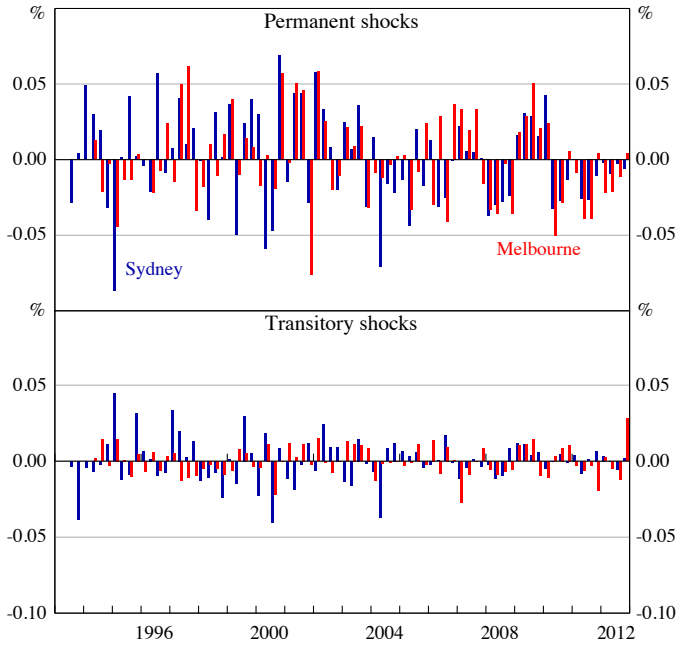

Figure 3. Estimates of Permanent and Temporary Shocks

In contrast, the temporary shocks are smaller. The most prominent periods of positive temporary shocks are the recovery from the early 1990s recession and around 2001 to 2003. During the global financial crisis, in contrast, there are no large temporary shocks – surprisingly, as for other countries the crisis has generally been interpreted as a demand shock and conventional wisdom is that macro demand

[23] See, for example, Ellis (2006) and Yates (2011).

shocks have only temporary effects on housing prices.

The impulse response functions to these shocks highlight two key properties (Figure 4). First, auction prices adjust more quickly than private-treaty prices to permanent price shocks. Around 80 (Sydney) to 60 (Melbourne) per cent of the long-run increase in auction prices occurs within the first quarter, but only around 35 to 25 per cent for private-treaty prices. After four quarters, roughly 95 per cent of the adjustment to the long-run auction price has occurred for both Sydney and Melbourne, but only 85 to 75 per cent for private-treaty prices. Second, temporary price shocks have smaller effects on private-treaty prices than do permanent shocks.

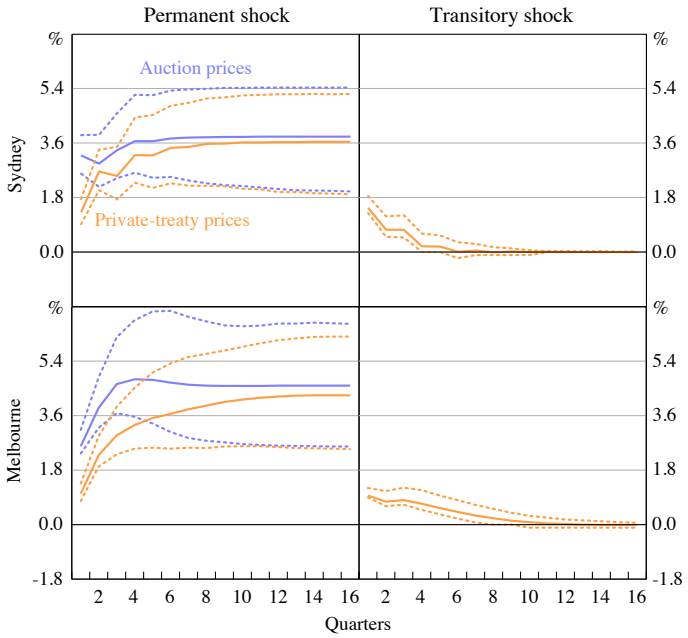

FIGURE 4. IMPULSE RESPONSE FUNCTIONS TO A ONE STANDARD DEVIATION SHOCK

Note: Impulse response functions to a one standard deviation shock estimated under the assumption that auction and private-treaty prices admit a VECM and that private-treaty prices do not Granger cause auction prices; confidence intervals are at the 95 per cent level of significance and are bootstrapped using Hall's percentile method.

That auction prices update quickly in response to permanent shocks is consistent with auctions being informational efficient. In contrast, the fact that private-treaty prices update more slowly, and are perturbed by temporary shocks, makes them less efficient this sense.

Are these findings robust? In the Appendix (Section VI.C) we explore four important dimensions in which they are. The first is measurement of the underlying

price indices. Using alternative hedonic or repeat-sales prices measures has little effect on our results (Table 14). Second, we account for endogenous selection of the sales mechanism by sellers. Using additional controls for auction incidence and the clearance rate to account for selection does not change the previous causality findings and selection appears to lag price dynamics rather than lead them (Table 16). Adjusting the underlying price indices for endogenous selection using a Heckman style structural model (Gatzlaff and Haurin (1997)) in which the past sale mechanism is allowed to determine the current propensity to use an auction but not the current price directly, leads to similar conclusions (Table 18). The Appendix also shows that our results are essentially unchanged when prices are adjusted for inflation (Table 20), which is unsurprising given that in most of our period of analysis Australia had inflation targeting that resulted in low and stable inflation.[24]

Finally, we consider within-group price variation. Looking at price within local sub-markets (districts), auctions continue to be more informative about local price trends than private-treaties (Table 21). The same is true conditioning on the type of home sold (Table 23). Together, these checks suggest our findings reflect differences in price formation by mechanism of sale, rather than more general housing market conditions that may be correlated with it.

IV. Interpreting the Results through Theory

We now examine a number of theoretical models aimed at interpreting the previous findings. All models are grounded in the micro structure of price formation. All conceive of prices as jointly determined by a pair of distributions of sellers' and buyers' valuations that evolve over time. The models differ in their predictions for how these distributions change in the short run, and in how auctions and negotiations map shifts in the distributions into price changes. Since the two price indices are cointegrated, the locations of these two distributions must follow the same stochastic trend in the long run. We assess these models according to both auxiliary data and variates of a small estimated state space model.

A. The Preferred Explanation: Asymmetric Weighting of Buyers' and Sellers' Valuations

[24] The literature on momentum considers both nominal prices (Titman, Wang, and Yang (2014)) and real prices (e.g. Case and Shiller (1989)).

Our preferred explanation is that, relative to private-treaty prices, auction prices are more responsive to buyers' than to sellers' shocks, and sellers' valuations lag buyers'. The first element is immediately evident when comparing the continuously ascending bid auction and the Nash bargaining solution. In the former, with which we model the English auction used in Australian housing markets, price equals the second highest bidder valuation. In the latter, price is a weighted average of the buyer and seller valuations.[25] Thus at auctions, a common shock to all bidders' valuations will increase the winning bid one for one. In negotiations, price will increase by less, according to the weight placed on the buyer's valuation (i.e., the seller's bargaining power).

The above assumes no seller reserve at auction and no overlap of the buyer and seller distributions. How their presence alters our claim depends on how the reserve price is formed and the particulars of the distributions. These issues are explored in the Appendix (Section VI.A). Here, we look at a simple example which makes clear that accounting for failed transactions does not fundamentally change the result that sellers valuations play less of a role in determining auction prices than they do private-treaty prices – a result that holds more generally. We assume the buyer and seller distributions are uniform, on $\left[\kappa^b, 1+\kappa^b\right]$ and $[\kappa^s, 1+\kappa^s]$, respectively, and a reserve price, set non-strategically equal to the seller's valuation, is announced to bidders at the auction's start. Then the average auction price is the expectation of the maximum of the second highest bidder valuation (denoted ν_2^b) and the seller reserve (ν^s), conditional on ν^s being less than the highest bidder's valuation (v_1^b) (otherwise, there is no sale), or

$$E\left(v_2^b | v_2^b \geq v^s\right) \frac{\Pr\left(v_2^b \geq v^s\right)}{\Pr\left(v_1^b \geq v^s\right)} + E\left(v^s | v_1^b \geq v^s > v_2^b\right) \frac{\Pr\left(v_1^b \geq v^s > v_2^b\right)}{\Pr\left(v_1^b \geq v^s\right)}$$

which can also be viewed as a weighted average of the second highest buyer's and seller's valuations, except that the weights are endogenous. The key point is that with sufficiently many bidders in auctions, typically six or more is enough, and with sufficient overlap in the distributions of buyers' and sellers' valuations to

[25] The Nash bargain per se is not important for the analysis. What matters is that bilateral trade prices reflect both the seller's and buyer's valuation. This is generally true in bargaining models with either complete or two-sided incomplete information (see for example, Myerson (1984) and Ausubel, Cramton and Deneckere (2002)). It is also consistent with seller price posting models in Caplin and Leahy (2011) and Díaz and Jerez (2013). What is important is that sellers' valuations are relatively more important in determining negotiated prices than they are auction prices.

match observed clearance rates, the probability with which the sellers' valuation determines the price outcome is small and insensitive to changes in either of the supports (κ^b or k^s). Indeed, it is considerably smaller than the probability with which price is determined by a buyers valuation.

Figure 5 makes this point clear, graphing the auction price and clearance rate as a function of κ^b (κ^s) in the left (right) panel, with κ^s set equal to 0.25 ($\kappa^b = 0$), and for six bidders.[26] The baseline choice of $\kappa^s - \kappa^b = 0.25$ yields a clearance rate that matches the average observed rate in the two cities. The figure shows that, within the range of auction clearance rates consistent with the data (the minimum and maximum observed across the two cities indicated by the red dashed lines and the average by the black dashed line), the auction price moves nearly one for one with perturbations to the distribution of buyers' valuations ($\Delta\kappa^b$), but is little changed with respect to perturbations in the seller's valuation distribution ($\Delta\kappa^s$). In contrast, the expected private-treaty price under equal bargaining power equals $\left(1 + k^b + k^s\right)/2$, so that the price is still equally affected by buyer and seller shocks when buyer and seller distributions overlap.[27] Section IV.D finds that bargaining power is equal in Sydney and insignificantly different from equality in Melbourne.

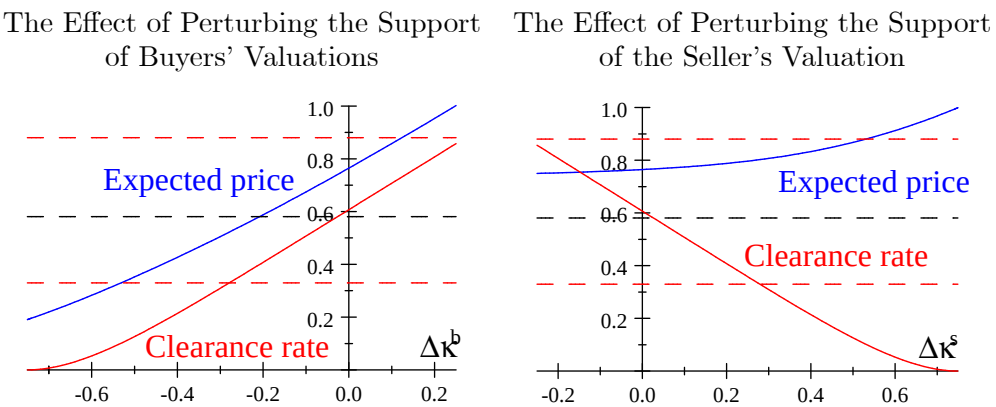

FIGURE 5. ASYMMETRY IN THE RESPONSE OF PRICE

[26] At higher bidder numbers, the results are even starker. For two bidders, shocks to the seller valuations have substantial effects, but they are still half as large as those for buyer shocks. We have no data source for the number of bidders, but newspaper articles report numbers that range from as low as one to as high as 45. Six seems typical, if somewhat on the low side.

[27] More generally, the expected private-treaty price is

$$(1-\psi)\,E\left(v^b|v^b \geq v^s\right) + \psi E\left(v^s|v^b \geq v^s\right) = \begin{cases} \frac{1}{3}\kappa^s + \frac{2}{3}k^b - \frac{1}{3}\psi + \frac{1}{3}\kappa^s\psi - \frac{1}{3}k^b\psi + \frac{2}{3} & \text{if } \kappa^s \geq k^b \\ -\frac{3a-6\kappa^b+\psi-3\kappa^s\psi+3\kappa^b\psi+3(\kappa^s)^2-3\left(\kappa^b\right)^2-2}{6\kappa^b-6\kappa^s+3} & \text{if } \kappa^s < k^b \end{cases}$$

where ψ is the weight on the seller's valuation – the buyer's bargaining power.

The second element in our preferred explanation is the assumption that sellers' valuations lags buyers'. A number of previously documented phenomena, such as the dramatic fall in seller time on the market and increase in the transaction price to list price ratio when demand increases (Genesove and Han (2012) and Haurin et al. (2013)), are consistent with that assumption. We provide additional suggestive evidence for the seller lag below.

In such an environment, and under a unit root process for buyers' valuations, prices behave qualitatively according to our empirical findings: auction prices Granger cause private-treaty prices, but not vice versa, the two prices are cointegrated, auction prices follow a random walk and there is positive momentum in private-treaty prices but not auction prices.

More formally, consider unit root process $z_t = \mu + z_{t-1} + \eta_t$, with η_t white noise, and let this be the common component of buyers' and sellers' valuations, so that auctions and private-treaty prices share the same stochastic trend. Buyer and seller valuations differ in that buyers' valuations capture all information in the common trend (z_t) contemporaneously, while sellers' only do so with a lag ($(1-\alpha) z_t + \alpha z_{t-1}$). Then, auction and private-treaty prices are given by

$$a_t = z_t \tag{8}$$

$$p_t = (1-\alpha\psi) z_t + \alpha\psi z_{t-1} \tag{9}$$

ψ is the weight put on the seller's valuation; it equals the buyer's surplus share.

This system generates all the time series properties documented earlier including: $a_t - p_t = -\alpha\psi\eta_t$, so private-treaty and auction prices are cointegrated with VECM representation $\Delta a_t = \eta_t, \Delta p_t = (a_{t-1} - p_{t-1}) + (1-\alpha\psi)\eta_t$; the two series admit a VAR (in levels) with $E[a_t \mid a_{t-1}, p_{t-1}] = E[p_t \mid a_{t-1}, p_{t-1}] = a_{t-1}$, so that auction prices Granger cause private-treaty prices but not vice-versa; and private-treaty prices display momentum, $Cov(\Delta p_t, \Delta p_{t-1}) = \psi\alpha(1-\alpha\psi) Var(\eta_t)$, while auction prices do not, $Cov(\Delta a_t, \Delta a_{t-1}) = 0$.

So far, the model lacks temporary shocks. Although we find that these shocks are unimportant for auctions prices,[28] they are a non-negligible part of private-treaty

[28] Note that we cannot reject the hypothesis that all shocks to auction prices are permanent, for both Sydney and Melbourne, in-sample.

price shocks. It is straightforward to incorporate such shocks by adding a stationary autoregressive moving average (ARMA) process to Equation (9). The end result in terms of the VECM representation, which continues to hold approximately, is, first, the addition of lagged growth terms in auction and private-treaty prices, and second, the coefficients on the cointegration term will now reflect the ARMA process as well as the basic parameters given above. Incorporating these temporary shocks thus frees up the specification from the strict cross equations restrictions in Equations (8) and (9), as observed empirically.

B. The Precision Explanation

One reason why temporary shocks may play a role in private-treaty price changes but not auction price changes is that temporary shocks may represent the time and mechanism specific aggregation of noisy signals around the 'true' value of search. The value of search is an important component of buyers' willingness to pay, and depends on expectations of future market conditions. If buyers have noisy signals of the true expectation, then there will be a common value component to their valuations. In that case, theory predicts conditioning on an individual auction price will provide a more precise prediction for future market conditions, and thus future prices, than conditioning on an individual private-treaty price.

The English auction – the standard mechanism for housing auctions in Australia – is particularly good at aggregating information. When there is a common component in bidders' valuations, and bidding strategies can condition on the exit of other bidders from the auction process, the auction price incorporates information from every buyer who actively makes a bid.[29] In contrast, prices determined through negotiation between a single buyer and seller incorporate information from those two parties only. Thus, an auction price may be a much less noisy predictor of future prices than a private-treaty price.

This argument requires that auction prices be less dispersed than private-treaty prices. In fact, as Table 7 shows, the root mean squared error (RMSE) of the hedonic regressions that underlie the price indices are similar for the two mechanisms. For Sydney, the RMSE is actually higher for auction prices. Furthermore, as we saw in Table 1, there are many more private-treaty transactions than auction trans-

[29] See Appendix D in Klemperer (1999) for example. This is more generally true in models with affiliated buyers' values (see Milgrom and Weber (1982)).

actions – ten times more in Sydney and six times more in Melbourne; consequently, the standard errors on the price indices are about 3 times larger for auctions than for private-treaties in Sydney, and twice as large for Melbourne. Indeed, even for auctions the number of transactions per quarter is so large that the contribution of transaction level variance to the variance of quarterly growth must be minimal, as comparing the standard deviation of quarterly price growth to the ratio of the RMSE to the square root of the average number of underlying observations in Table 7 (columns three and four) shows. Similar results obtain for repeat-sales regressions considered in the Appendix (Section VI.C, Table 14).

Table 7—RMSE of Hedonic Price Regressions

	RMSE	N	$\frac{RMSE}{\sqrt{N}}$	St. Dev.
		Sydney		
Auction	0.27	911	0.009	0.034
Private	0.24	5 034	0.003	0.031
		Melbourne		
Auction	0.34	2 604	0.007	0.030
Private	0.36	16 128	0.003	0.022

Note: RMSE is the root mean squared error of the corresponding regression; N is the average number of observations per quarter; and St. Dev. is the standard deviation of quarterly prices growth.

Finding near equal RMSEs for the two mechanisms should not be interpreted as a rejection of common value auction theory or indicating the absence of a common value component. Other factors, such as the variance of unobserved quality, also contribute to the RMSE. However, along with the comments on the number of observations, it does indicate that any explanation of our findings based on temporary shocks cannot be sourced at the individual transaction level.

An unequal variance in the temporary shocks might still explain why the two price indices differ in their predictive ability, if those shocks are common to many underlying transactions and are not eliminated by aggregation. One possible source of common temporary shocks is changing bargaining weights. Although some bargaining may take place after a winning bid is rejected at auction, bargaining is not an integral part of the auction process; shocks to bargaining weights could thus explain why temporary shocks are so much more important for private-treaty prices than auction prices. Volatility in bargaining weights does not arise naturally in the Nash bargaining solution, but could arise in other solutions and in environments with changing private information, as suggested by Kennan (2010).

C. Kalman Filter Estimates

Equations (8)–(9) are easily amended to accommodate the above explanation, in which auction and private-treaty averages are noisy indicators of permanent common shocks. Combining our preferred model with this precision model yields

$$a_t = \beta z_t + \varepsilon_t^a \tag{10}$$

$$p_t = (1 - \alpha\psi)\, z_t + \alpha\psi z_{t-1} + \varepsilon_t^p \tag{11}$$

where ε_t^a and ε_t^P are each white noise. β is added to account for the non-unitary coefficient in the error correction term documented earlier. The price indices remain cointegrated in this extended model; the remaining qualitative properties of (8)–(9) continue to hold if $Var\,(\varepsilon_t^a)$ is small. Obviously α and ψ are not separately identified. A necessary condition for the precision explanation to be valid is $Var\,(\varepsilon_t^a) \leq Var\,(\varepsilon_t^p)$. The preferred model corresponds to $0 < \alpha\psi < 1$.

Tables 8 and 9 present Kalman Filter estimates of model (10)–(11) for Sydney and Melbourne, respectively, under various restrictions and generalizations.[30] Column (1) shows estimates of the basic model. They are much more in line with the preferred than with the precision explanation. On the one hand, with $\widehat{\alpha\psi}$ equal to 0.52 in Sydney and 0.70 in Melbourne, the private-treaty price puts about half (seventy percent) of its weight on the lagged state variable in Sydney (Melbourne). On the other hand, the precision model does very poorly. In Sydney, the variances of the temporary shocks are very similar, and one cannot reject equality between them. The failure of the precision model for Melbourne is even starker, with the variance of the auction temporary shock about twenty times larger that of the private-treaty shock. The remaining columns show that Column (1)'s restrictions on the lags – no lag for the auction price and one (two) lags for the Sydney (Melbourne) private-treaty price – are not rejected by the data.

D. Evidence from List Prices

List prices are set solely by sellers and so should reflect seller information only.[31] Incorporating list prices into the preferred model we then have that auction prices

[30] Here, and in the tables that follow, we typically use zero restrictions on the correlations between shocks to help identify them. However, where identification is still achieved without those restrictions, we relax them and report the estimated correlation coefficients.

[31] This could include sellers' beliefs about buyers' valuations, but not uncorrelated contemporaneous shifts in the actual buyer distribution.

Table 8—Structural Unobserved Componsents Models – Sydney

Coefficient	(1)	(2)	(3)
A: z_t	1	1	0.79***
	(.)	(.)	(0.14)
A: z_{t-1}			0.21
			(0.14)
P: z_t	0.44***	0.45***	0.21
	(0.10)	(0.10)	(0.14)
P: z_{t-1}	0.52***	0.48***	0.75***
	(0.10)	(0.15)	(0.14)
P: z_{t-2}		0.03	
		(0.09)	
σ_η^2	1.17***	1.21***	1.16***
	(0.27)	(0.31)	(0.27)
σ_a^2	0.18	0.16	0.44***
	(0.11)	(0.14)	(0.13)
σ_p^2	0.22***	0.22***	0.13
	(0.06)	(0.06)	(0.11)
$corr(\varepsilon_t^a, \varepsilon_t^p)$	-0.54	-0.64	
	(0.39)	(0.58)	
	Associated p-values		
$H_0 : \sigma_a^2 = \sigma_p^2$	0.81	0.70	0.16
$H_0 : \sigma_a^2 = \sigma_{ap} = 0$	0.00***	0.01***	
Log Likelihood	363.19	363.44	367.66

Note: ***, ** and * denote significance at the 1, 5 and 10 per cent levels. Standard errors in parentheses. Variance estimates and their standard errors are multiplied by 1000. σ_η^2 is the variance of the permanent shock, η_t. σ_a^2 and σ_p^2 are the variances of the temporary shocks to auction and private-treaty prices respectively, $corr\left(\varepsilon_t^a, \varepsilon_t^p\right)$ their correlation and σ_{ap} their covariance.

reflect buyer information only, private-treaty prices reflect both buyer and seller information, while list prices reflect seller information only; also, private-treaty prices lag auction prices, and list prices lag private-treaty prices.

We formed a list price index in the same manner used for the other indices, assigning a property to its first quarter of listing. There being few auctions for which a list price is available, we only examine list prices for homes offered for sale via private treaty and with a recorded list price.[32] Lacking list prices for sales prior to 1998:II, our sample size drops to 58 observations only.[33]

We first run Granger causality tests for list prices and the two other series (Table 10). Our Sydney results are perfectly in line with the preferred model: both auction prices and private-treaty prices Granger cause list prices, but list prices

[32] Including list prices for auctions has little overall effect on the results. We do not have information on properties offered for sale by private-treaty that were withdrawn from the market.

[33] Nevertheless, the shorter sample allows us to use hedonic indices with all attributes data.

Table 9—Structural Unobserved Componsents Models – Melbourne

Coefficient	(1)	(2)	(3)
A: z_t	1	1	0.81***
	(.)	(.)	(0.21)
A: z_{t-1}			0.28
			(0.21)
P: z_t	0.22	0.51***	0.41***
	(0.15)	(0.07)	(0.12)
P: z_{t-1}	0.70***	0.12*	0.19***
	(0.15)	(0.07)	(0.07)
P: z_{t-2}		0.29***	0.40***
		(0.05)	(0.10)
σ_η^2	0.67***	1.03***	0.95***
	(0.14)	(0.21)	(0.20)
σ_a^2	1.04***	0.89***	0.87***
	(0.36)	(0.16)	(0.15)
σ_p^2	0.05	0.03	0.04***
	(0.06)	(0.02)	(0.01)
$corr(\varepsilon_t^a, \varepsilon_t^p)$	0.29		
	(0.78)		
	Associated p-values		
$H_0 : \sigma_a^2 = \sigma_p^2$	0.01***	0.00***	0.00***
$H_0 : \sigma_a^2 = \sigma_{ap} = 0$	0.00***		
Log Likelihood	348.16	360.25	360.93

Note: ***, ** and * denote significance at the 1, 5 and 10 per cent levels. Standard errors in parentheses. Variance estimates and their standard errors are multiplied by 1000. Melbourne data are seasonally adjusted prior to estimation. σ_η^2 is the variance of the permanent shock, η_t. σ_a^2 and σ_p^2 are the variances of the temporary shocks to auction and private-treaty prices respectively, $corr\left(\varepsilon_t^a, \varepsilon_t^p\right)$ their correlation and σ_{ap} their covariance.

Granger causes neither of them. The second statement holds for Melbourne as well. However, there, only auction prices Granger causes list prices.

Expanding the state space model to include a list price index, l_t, allows us to separately identify α and ψ under the model

$$a_t = \beta z_t + \varepsilon_t^a \tag{12}$$

$$p_t = (1 - \alpha\psi)\, z_t + \alpha\psi z_{t-1} + \varepsilon_t^p \tag{13}$$

$$l_t = (1 - \alpha)\, z_t + \alpha z_{t-1} + \varepsilon_t^l \tag{14}$$

Table 11, Columns (1) and (2), presents Kalman Filter estimates of this model. Although the samples are shorter, the result that private-treaty prices lag the cycle continues to hold in both cities. Where precisely estimated, in Sydney, the bargaining weight on the seller's valuation, ψ, is estimated at 0.5 (equal bargaining

Table 10—Granger Causality Results Including List Prices

Null Hypothesis	Sydney	Melbourne
H_0 : List prices do not	3.36	3.73
Granger Cause auction prices	(0.50)	(0.44)
H_0 : List prices do not	1.47	3.64
Granger Cause private-treaty prices	(0.83)	(0.45)
H_0 : Auction prices do not	21.20***	12.36**
Granger Cause list prices	(0.00)	(0.01)
H_0 : Private-treaty prices do not	10.32**	5.10
Granger Cause list prices	(0.04)	(0.27)

Note: ***, ** and * denote significance at the 1, 5 and 10 per cent levels of significance; test statistics constructed using the approach outlined in Toda and Yamamoto (1995); p-values in parentheses.

power) and the backward looking component in sellers' valuations, α, at 0.83. As $\psi = 0$ ($\psi = 1$) would imply that private-treaty prices have the same autocorrelation properties as auction prices (list prices), the data do not clearly favour either boundary case: private-treaties prices are consistent with a convex combination of both buyers' and sellers' values. For Melbourne, the respective estimates are 0.82 – insignificantly different from both equal and zero seller bargaining power[34] – and 0.31. Overall, the behaviour of the list price index accords with our preferred model and allows us to identify the bargaining weight.

E. Evidence from Auction Clearance Rates

One of the notable stylised facts from auctions is the strong contemporaneous correlation between auction prices and the auction clearance rate. Figure 6 is a scatter plot of quarterly growth in prices and the quarterly clearance rate, superimposed by the line of best fit. The contemporaneous correlations between price growth and the clearance rate are 0.37 in Sydney and 0.40 in Melbourne. An important question is whether the preferred model is also able to replicate this fact.

With sellers' valuations lagging buyers', innovations to a common stochastic trend will tend to affect buyers bids earlier than sellers' reserve prices. This should generate positive correlation between price growth and the clearance rate. To see this, let a buyer's valuation of a property at time t be $z_t + b$, where z_t is again the common component of buyers' and sellers' valuations, while b is specific to the buyer-property match and is drawn from some distribution. Likewise, let the

[34] Price posting, as in Díaz and Jerez (2013) and Caplin and Leahy (2011), is an alternative interpretation of $\psi = 1$.

Table 11—Unobserved Components Models with Listing Prices

Parameter	Sydney (1)	Melbourne (2)	Sydney (3)	Melbourne (4)
β	1.02***	1.06***	0.98***	1.06***
	(0.03)	(0.01)	(0.06)	(0.01)
α or $\delta^{\dagger}$	0.83***	0.31**	0.75***	0.13
	(0.23)	(0.14)	(0.14)	(0.12)
α_2 or $\delta_2^{\dagger}$		0.17**		0.08**
		(0.10)		(0.05)
ψ	0.50***	0.82***	0.54***	0.87**
	(0.13)	(0.23)	(0.09)	(0.53)
σ_η^2	0.54***	0.98***	0.87***	0.79***
	(0.13)	(0.22)	(0.21)	(0.18)
σ_a^2	0.66***	0.43***		0.64***
	(0.20)	(0.13)		(0.17)
σ_p^2	0.25***	0.18***	0.28***	0.13***
	(0.08)	(0.06)	(0.05)	(0.06)
σ_l^2	0.08**	0.86***	0.85***	0.90***
	(0.09)	(0.18)	(0.16)	(0.19)
$corr\left(\varepsilon_t^a, \varepsilon_t^p\right)$	0.71***			
	(0.12)			
$corr\left(\varepsilon_t^a, \varepsilon_t^l\right)$	-0.55***			
	(0.17)			
$corr\left(\varepsilon_t^p, \varepsilon_t^l\right)$	-0.87**		0.58***	
	(0.47)		(0.09)	
μ	1.35***	1.93***	1.35	1.92***
	(0.31)	(0.41)	(2.72)	(0.37)
H_0 : No	p-values		p-values	
lagged diffusion††	0.00	0.00	0.00	0.03
Log Likelihood	425.33	364.16	401.07	358.08

Note: ***, ** and * denote significance at the 1, 5 and 10 per cent levels of significance. Variances estimates and their standard errors are multiplied by 1000; Melbourne data are seasonally adjusted prior to estimation and include an additional lag for the diffusion of common shocks. $^{\dagger}\alpha$ refers to columns one and two, δ refers to columns three and four. ††No lagged diffusion denotes $H_0 : \alpha = 0$ ($\delta = 0$) for Sydney and $H_0 : \alpha = \alpha_2 = 0$ ($\delta = \delta_2 = 0$) for Melbourne.

seller's valuation of a property in the market at time t be $(1-\alpha)\, z_t + \alpha z_{t-1} + s$, where s is specific to the seller and is drawn from some other distribution. Then the probability of sale is

$$\Pr\left(b^{(1)} - s \geq -\alpha\eta_t\right) \equiv h\left(\alpha\eta_t\right)$$

where $b^{(j)}$ is the jth order statistic of b ($b^{(1)} \geq ... \geq b^{(N)}$), and the expected auction price is

$$\begin{aligned} a_t &= z_t + E_{b,s:N}\left[\max\left(b^{(2)}, s\right) \,\middle|\, b^{(1)} \geq s - \alpha\eta_t\right] \\ &\approx z_t - \alpha q^a \eta_t \end{aligned}$$

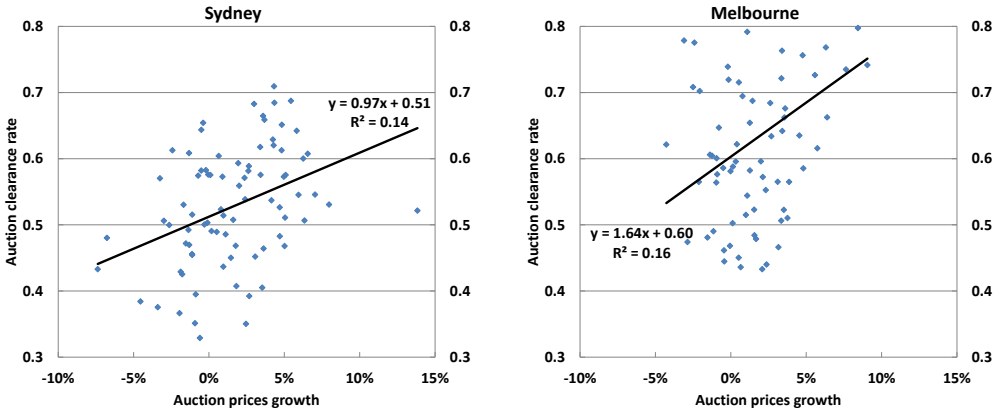

FIGURE 6. SCATTER PLOT OF AUCTION PRICE GROWTH AGAINST THE CLEARANCE RATE

with the probability and expectation taken with respect to the joint distribution of s and N draws of b, and q^a measures the extent to which the expected auction price changes due to the presence of a reserve price

$$q^a \equiv \frac{\partial E_{b,s:N}\left[\max\left(b^{(2)}, s\right) \,\middle|\, b^{(1)} - s \geq x\right]}{\partial x} \text{ evaluated at } x = 0.$$

The correlation between auction prices and the probability of sale is given by

$$\begin{aligned} Cov\left(\Delta a_t, h\left(\alpha\eta_t\right)\right) &\approx \left(1 - \alpha q^a\right)\alpha h'\left(0\right)\sigma_\eta^2 \\ &> 0 \text{ if } \alpha q^a < 1 \end{aligned}$$

Thus, provided the interaction of lagged information diffusion and the clearance effect on auction prices, αq^a, is not too large, auction prices and the clearance rate will be positively correlated. The calculations in the Appendix (Section VI.A) show that $q^a < 1$ for all possible pairs of the Generalised Pareto distributions, with $2 \leq N \leq 30$, differences in support that match observed clearance rates, and for alternative assumptions about the reserve price.

F. *Differential Weighting of the Buyer Valuation Distribution*

Auctions and negotiations differ in the weight that each gives to different parts of the buyer distribution. If shocks diffuse through the buyer population over time, this can affect the lead-lag relationship between auction and private-treaty prices. Whether or not the resulting pattern can fit what we observe depends on the nature

of information that buyers have, and accumulate, at the auction.

Under private values, positive shocks to valuations of a fraction of the buyer population are felt more in auction prices, while negative shocks are felt more in negotiations, given a sufficient number of bidders at auction. For a positive shock, those receiving it tend to outbid the remaining buyers, and so price reflects the shock; when negative, those receiving the shock are outbid and price does not reflect it. In negotiations, in contrast, price reflects the shock, regardless of sign, whenever a "shocked" buyer is present. This reasoning suggests that auction prices will lead private-treaty prices when shocks are positive, but lag when negative. As usual, unconsummated sales blur the distinction, but the general claim that the right tail of the buyer distribution is relatively more important in auctions presumably continues to hold.

For a simple example, let a random fraction a of buyers receive a positive shock in the first period, and $1-a$ in the second. Buyers' valuations are identical prior to the shocks. Then price at any given auction increases by the amount of the shock if at least two bidders there have received it; the expected price at auction increases, per unit of the shock, by $q(a) \equiv 1-(1-a)^N - N(1-a)^{N-1} a$, and by the remaining $1-q(a)$ in the next period. In private treaties, price increases in the first period so long as the buyer has received the shock, and zero otherwise. Percentage-wise, then, auction prices increase more than private-treaty prices so long as $q(a) > a$, which holds for $a \in (a^*(N), 1)$, where a^* is a declining function of N. For example, $a^*(4) = 0.24$ and $a^*(8) = 0.04$. In contrast, for a negative shock, the auction price falls only if all or all but one of the bidders have received it, so that the expected decrease is $1-q(1-a)$. Percentage-wise, auction prices fall less than private-treaty prices so long as $1-q(1-a) < a$, which holds for $a \in (1-a^*(N), 1)$.

In principle, this mechanism could explain our results if the auction-leading-private treaty effect is larger than the converse. In the Appendix (Section VI.B), we conduct a gross check on the explanation by looking at the lagged cross correlation of auction and private-treaty price growth. If this explanation is correct, auction price growth should be positively correlated with one-period-ahead growth in private-treaty prices when auction prices are increasing, and private-treaty price growth positively correlated with one period ahead growth in auction prices when

private-treaty prices are falling. We find the explanation inconsistent with the data, as private-treaty price growth does not lead auction price growth when private-treaty prices are falling, or rising less than usual. As an alternative structural check, we estimate non-linear models that allow for contemporaneous and lagged asymmetry in the response of auction prices that depends on the direction of change in the permanent component of prices. The results, also shown in the Appendix (Section VI.B, Table 13), find no evidence of asymmetry that would be consistent with lagged diffusion of shocks to buyers.

Can common values, or affiliated values, rescue this argument? Such models are difficult to work with, and, to our knowledge, no one has analysed a model with a signal distribution that shifts over time. Our comments are thus impressionistic only. The crucial issue here is the extent to which the winning bid in ascending auctions incorporates all bidders' signals, or just particular order statistics. That, in turn, depends on what is observed at the auction. If bidders do not observe other bidders dropping out, then the auction is equivalent to a second price sealed bid auction, with price a function of the second order statistic. Then the same argument will hold as before, and auction prices will lead private-treaty prices when shocks are positive, and lag them when shocks are negative. However, if the bidders do observe exits, then all active bidders' signals matter, and so auction prices will in general be sensitive to both positive and negative shocks.

The signals' weights in price for the few cases that have been worked out depend on how valuations depend on the signals. For linear affiliated values, the second order statistic matters more than the other signals, which are weighted equally, which returns us to the private values case described above. In the uniform distribution case, the auction price equals the average signal plus the gap between the first order and second order bid statistics, divided by the number of bidders. For large numbers of bidders, the percentage change in price per additional unit of valuation will become close to a, the same as that for private-treaty prices. Thus, affiliated values coupled with slow diffusion of shocks across the buyer distribution, are unlikely to generate leading auction prices.

G. Backward Looking Price Formation and Publicity

According to this explanation, buyers and sellers rely on previous prices to form market valuations, weighting more salient prices more heavily. There are several

reasons why both buyers and sellers might form their valuations based on past transactions values: (a) appraised values, which rely on past transactions (Quan and Quigley (1991)), are publically available; (b) prices are revealing about the current state of the market (this is relevant given heterogeneity in homes sold and the absence of centralised markets for trade); and (c) expectations may be backward looking (Case and Shiller's surveys, 1988) or subject to information rigidities (Coibion and Gorodnichenko (2015)). Past auction prices are better publicised – they are more dramatic than negotiations, attended by more people and concurrently published in newspapers and on auction company websites – and thus have a greater effect on current prices. In contrast, the delay between reaching a private agreement and the publication of the resulting price can often be a quarter or longer.

Price formation that focuses on recent auction prices implies Granger causality from auction to private-treaty prices and momentum in the latter. However, it also generates equally large auction price momentum. To see this, write the model

$$a_t = \delta z_t + (1-\delta)\, a_{t-1} + \varepsilon_t^a \tag{15}$$

$$p_t = \delta z_t + (1-\delta)\, a_{t-1} + \varepsilon_t^p \tag{16}$$

This model asserts that prices are a convex combination of the current state and the lagged auction price, plus a temporary shock. We use the same coefficient in both equations under the assumption that the use of historical information is independent of the sale mechanism. Rewriting these equations, we obtain

$$\Delta a_t = \mu + \delta m(\delta)\, \eta_t + m(\delta)\, \Delta\varepsilon_t^a \tag{17}$$

$$\Delta p_t = \mu + \delta\left(1 + (1-\delta)\, m(\delta)\, L\right)\eta_t + (1-\delta)\, m(\delta)\, \Delta\varepsilon_{t-1}^a + \Delta\varepsilon_t^p \tag{18}$$

where $m(\delta) = (1-(1-\delta)\, L)^{-1}$. Equation (17) makes clear that δ must be close to 1 to be consistent with the auctions data. Otherwise, auction price growth is a linear combination of two (independent) infinite moving average prices and so could be approximated by a low-order autoregressive process – i.e. would be significantly autocorrelated.[35] However, for δ close to one, private-treaty price growth should

[35] To see this even more clearly, set $\varepsilon_t^a = 0$ and simply first difference Equation (15). The result is simply an AR(1) process in auction price growth whose persistence is decreasing in δ.

also lack autocorrelation. This is inconsistent with our evidence.

Can asymmetry in the use of historical information rescue this argument? In principle, yes. If only sellers condition their information on past auction prices, when forming valuations, we have the model

$$
\begin{aligned}
a_t &= \beta z_t + \varepsilon_t^a \\
p_t &= (1-\psi)\, z_t + (1-\delta)\, \psi z_t + \psi \delta a_{t-1} + \varepsilon_t^p \\
l_t &= (1-\delta)\, z_t + \delta a_{t-1} + \varepsilon_t^l
\end{aligned}
$$

The only substantive difference from the preferred model is that sellers condition on lagged auction prices rather than the unobserved permanent component itself. Assuming sellers use past auction prices is consistent with our previous findings. Estimates of this model are reported in Columns (3) and (4) of Table 11. The results are similar to the preferred model, including the estimates of relative bargaining strength and the weight on past auction prices.

V. Conclusion

Housing market dynamics differ dramatically from what perfect asset models predict and so have proved difficult to model successfully. Particularly challenging has been the high positive autocorrelation of housing price growth, which has been documented for a large number of cities, countries and time periods. Working in an environment with an unusually high share of auction sales, we find a much lower autocorrelation in auction prices than negotiated sales, which other markets use near exclusively. We argue that the larger weight that auction prices put on buyers' valuations rather than sellers', in comparison to negotiated prices, points to seller valuations as the source of the autocorrelation. We argue, further, that seller valuations appear to lag buyer valuations, and provide supporting evidence for this claim in the behaviour of list prices and the Phillips curve like relationship between the auction clearance rate and price growth.

Indeed, recent calibration studies have incorporated seller sluggishness in order to generate positive price growth autocorrelation. However, why sellers update values more slowly than buyers in response to new shocks is unclear. We suspect that the asymmetric nature of the matching process is responsible for the most part, for

the reasons given in the Introduction. These explanation need further theoretical elaboration, and empirical verification, which should advance our understanding of housing market dynamics even further. We hope that this paper can serve as an example of how the differential price dynamics, reflected in different trading mechanisms, can serve as a prism for underling market behaviour.

REFERENCES

Anenberg, Elliot. 2011. "Loss Aversion, Equity Constraints and Seller Behavior in the Real Estate Market." *Regional Science and Urban Economics*, 41(1): 67–76.

Ausubel, Lawrence M., Peter Cramton, and Raymond J. Deneckere. 2002. "Bargaining with Incomplete Information." In *Handbook of Game Theory with Economic Applications*. Vol. 3, ed. Robert J. Aumann and Sergiu Hart, 1897–1945. Elsevier.

Berkovec, James A., and John L. Goodman. 1996. "Turnover as a Measure of Demand for Existing Homes." *Real Estate Economics*, 24(4): 421–440.

Bulow, Jeremy, and Paul Klemperer. 1996. "Auctions Versus Negotiations." *American Economic Review*, 86(1): pp. 180–194.

Bulow, Jeremy, and Paul Klemperer. 2009. "Why Do Sellers (Usually) Prefer Auctions?" *American Economic Review*, 99(4): 1544–75.

Caplin, Andrew, and John Leahy. 2011. "Trading Frictions and House Price Dynamics." *Journal of Money, Credit and Banking*, 43: 283–303.

Capozza, Dennis R., Patric H. Hendershott, and Charlotte Mack. 2004. "An Anatomy of Price Dynamics in Illiquid Markets: Analysis and Evidence from Local Housing Markets." *Real Estate Economics*, 32(1): 1–32.

Case, Karl, and Robert Shiller. 1988. "The Behavior of Home Buyers in Boom and Post-boom Markets." *New England Economic Review*, November: 29–46.

Case, Karl E., and Robert J. Shiller. 1989. "The Efficiency of the Market for Single-Family Homes." *American Economic Review*, 79(1): 125–137.

Case, Karl E., Robert J. Shiller, and John M. Quigley. 2005. "Comparing Wealth Effects: The Stock Market versus the Housing Market." *Advances in Macroeconomics*, 5(1): 1–34.

Cho, Man. 1996. "House Price Dynamics: A Survey of Theoretical and Empirical Issues." *Journal of Housing Research*, 7(2): 145–172.

Clark, Todd E., and Kenneth D. West. 2007. "Approximately Normal Tests for Equal Predictive Accuracy in Nested Models." *Journal of Econometrics*, 138(1): 291–311.

Coibion, Olivier, and Yuriy Gorodnichenko. 2015. "Information Rigidity and the Expectations Formation Process: A Simple Framework and New Facts." *American Economic Review*, 105(8): 2644–78.

Čopič, Jernej, and Clara Ponsatí. 2008. "Robust Bilateral Trade and Mediated Bargaining." *Journal of the European Economic Association*, 6(2-3): 570–580.

Cross, John G. 1969. *The Economics of Bargaining.* Basic Books New York.

Cutler, David M., James M. Poterba, and Lawrence H. Summers. 1991. "Speculative Dynamics." *The Review of Economic Studies*, 58(3): 529–546.

Díaz, Antonia, and Belén Jerez. 2013. "House Prices, Sales, And Time on the Market: A Search-Theoretic Framework." *International Economic Review*, 54(3): 837–872.

Diebold, Francis X, and Roberto S Mariano. 1995. "Comparing Predictive Accuracy." *Journal of Business & Economic Statistics*, 13(3): 253–263.

Dumitrescu, Elena-Ivona, and Christophe Hurlin. 2012. "Testing for Granger Non-causality in Heterogeneous Panels." *Economic Modelling*, 29(4): 1450 – 1460.

Einav, Liran, Theresa Kuchler, Jonathan Levin, and Neel Sundaresan. 2015. "Assessing Sale Strategies in Online Markets Using Matched Listings." *American Economic Journal: Microeconomics*, 7(2): 215–47.

Ellis, Luci. 2006. "Housing and Housing Finance: The View from Australia and Beyond." RBA Research Discussion Paper 2006–12.

Engelhardt, Gary V. 2003. "Nominal Loss Aversion, Housing Equity Constraints, and Household Mobility: Evidence from the United States." *Journal of Urban Economics*, 53(1): 171–195.

Fisher, Lance A., and Hyeon-Seung Huh. 2007. "Permanent-Transitory Decompositions under Weak Exogeneity." *Econometric Theory*, 23(1): 183–189.

Fudenberg, D., and J. Tirole. 1991. *Game Theory.* MIT Press.

Gatzlaff, Dean H., and Donald R. Haurin. 1997. "Sample Selection Bias and Repeat-Sales Index Estimates." *The Journal of Real Estate Finance and Economics*, 14(1): 33–50.

Genesove, David, and Christopher Mayer. 1997. "Equity and Time to Sale in the Real Estate Market." *American Economic Review*, 87(3): 255–269.

Genesove, David, and Christopher Mayer. 2001. "Loss Aversion and Seller Behavior: Evidence from the Housing Market." *The Quarterly Journal of Economics*, 116(4): 1233–1260.

Genesove, David, and Lu Han. 2012. "Search and Matching in the Housing Market." *Journal of Urban Economics*, 72(1): 31–45.

Glaeser, Edward L., Joseph Gyourko, Eduardo Morales, and Charles G. Nathanson. 2014. "Housing Dynamics: An Urban Approach." *Journal of Urban Economics*, 81: 45–56.

Guren, Adam M. 2015. "The Causes and Consequences of House Price Momentum." Harvard University Mimeo.

Han, Lu, and William C. Strange. 2014. "Bidding Wars for Houses." *Real Estate Economics*, 42(1): 1–32.

Hansen, James. 2009. "Australian House Prices: A Comparison of Hedonic and Repeat-Sales Measures." *Economic Record*, 85(269): 132–145.

Harsanyi, John C. 1956. "Approaches to the Bargaining Problem Before and After the Theory of Games: A Critical Discussion of Zeuthen's, Hicks', and Nash's Theories." *Econometrica*, 24(2): pp. 144–157.

Haurin, Donald, Stanley McGreal, Alastair Adair, Louise Brown, and James R. Webb. 2013. "List Price and Sales Prices of Residential Properties During Booms and Busts." *Journal of Housing Economics*, 22(1): 1–10.

Head, Allen, Huw Lloyd-Ellis, and Hongfei Sun. 2014. "Search, Liquidity, and the Dynamics of House Prices and Construction." *American Economic Review*, 104(4): 1172–1210.

Head, Allen, Huw Lloyd-Ellis, and Hongfei Sun. 2016. "Search, Liquidity, and the Dynamics of House Prices and Construction: Corrigendum." *American Economic Review*, 106(4): 1214–19.

Iacoviello, Matteo. 2005. "House Prices, Borrowing Constraints, and Monetary Policy in the Business Cycle." *American Economic Review*, 95(3): 739–764.

Iacoviello, Matteo, and Stefano Neri. 2010. "Housing Market Spillovers: Evidence from an Estimated DSGE Model." *American Economic Journal: Macroeconomics*, 2(2): 125–64.

Inoue, Atsushi, and Lutz Kilian. 2005. "In-Sample or Out-Of-Sample Tests of Predictability: Which One Should We Use?" *Econometric Reviews*, 23(4): 371–402.

Kennan, John. 2010. "Private Information, Wage Bargaining and Employment Fluctuations." *The Review of Economic Studies*, 77(2): 633–664.

Klemperer, Paul. 1999. "Auction Theory: A Guide to the Literature." *Journal of Economic Surveys*, 13(3): 227–286.

Krainer, John. 2001. "A Theory of Liquidity in Residential Real Estate Markets." *Journal of Urban Economics*, 49(1): 32–53.

Kremer, Ilan. 2002. "Information Aggregation in Common Value Auctions." *Econometrica*, 70(4): 1675–1682.

Leamer, Edward E. 2007. "Housing is the Business Cycle." Federal Reserve Bank of Kansas City Proceedings – Economic Policy Symposium – Jackson Hole.

Lucking-Reiley, David. 1999. "Using Field Experiments to Test Equivalence between Auction Formats: Magic on the Internet." *American Economic Review*, 89(5): 1063–1080.

McCracken, Michael W. 2007. "Asymptotics for Out of Sample Tests of Granger Causality." *Journal of Econometrics*, 140(2): 719–752.

Merzyn, Wolfram, Gabor Virag, and Stephan Lauermann. 2010. "Aggregate Uncertainty and Learning in a Search Model." Society for Economic Dynamics 2010 Meeting Papers 1235.

Milgrom, Paul R., and Robert J. Weber. 1982. "A Theory of Auctions and Competitive Bidding." *Econometrica*, 50(5): 1089–1122.

Myerson, Roger B. 1984. "Two-Person Bargaining Problems with Incomplete Information." *Econometrica*, 52(2): pp. 461–488.

Napel, S. 2002. *Bilateral Bargaining: Theory and Applications. Lecture Notes in Economics and Mathematical Systems*, Springer Berlin Heidelberg.

Nash, John. 1950. "The Bargaining Problem." *Econometrica*, 18(2): 155–162.

Pagan, Adrian R., and M. Hashem Pesaran. 2008. "Econometric Analysis of Structural Systems with Permanent and Transitory Shocks." *Journal of Economic Dynamics and Control*, 32(10): 3376–3395.

Piazzesi, Monika, and Martin Schneider. 2009. "Momentum Traders in the Housing Market: Survey Evidence and a Search Model." *American Economic Review*, 99(2): 406–411.

Prasad, Nalini, and Anthony Richards. 2008. "Improving Median Housing Price Indexes through Stratification." *Journal of Real Estate Research*, 30(1): 45–72.

Quan, Daniel C., and John M. Quigley. 1991. "Price Formation and the Appraisal Function in Real Estate Markets." *Journal of Real Estate Finance and Economics*, 4(2): 127–146.

Roth, A.E. 2012. *Axiomatic Models of Bargaining. Lecture Notes in Economics and Mathematical Systems*, Springer Berlin Heidelberg.

Sommervoll, Dag Einar, Trond-Arne Borgersen, and Tom Wennemo. 2010. "Endogenous Housing Market Cycles." *Journal of Banking & Finance*, 34(3): 557–567.

Stein, Jeremy C. 1995. "Prices and Trading Volume in the Housing Market: A Model with Down-Payment Effects." *The Quarterly Journal of Economics*, 110(2): 379–406.

Titman, Sheridan, Ko Wang, and Jing Yang. 2014. "The Dynamics of Housing Prices." National Bureau of Economic Research Working Paper 20418.

Toda, Hiro Y., and Taku Yamamoto. 1995. "Statistical Inference in Vector Autoregressions with Possibly Integrated Processes." *Journal of Econometrics*, 66(1–2): 225–250.

Wang, Ruqu. 1993. "Auctions versus Posted-Price Selling." *American Economic Review*, 83(4): 838–851.

Wang, Ruqu. 1998. "Auctions versus Posted-Price Selling: The Case of Correlated Private Valuations." *The Canadian Journal of Economics/Revue canadienne d'économique*, 31(2): pp. 395–410.

Wheaton, William C. 1990. "Vacancy, Search, and Prices in a Housing Market Matching Model." *Journal of Political Economy*, 98(6): pp. 1270–1292.

Yates, Judith. 2011. "Housing in Australia in the 2000s: On the Agenda Too Late?" In *The Australian Economy in the 2000s. Proceedings of a Conference*, 261–296. Sydney: Reserve Bank of Australia.

VI. Appendix

A. Accounting for failed transactions

This section generalizes Section IV.A's discussion of how a seller reserve and overlapping buyer and seller distributions – which are necessary to account for not all meetings of buyers and sellers ending in a sale – affect the preferred model's ability to account for our core empirical results on Granger causality and momentum. We consider generalisations of the buyer and seller distributions and alternative models of the seller's reserve to those used in that section.

Let the valuation of a given home to a potential buyer in the market at time t be $z_t + b$ where z_t is common to all buyers, while b is specific to the buyer-home pair and is drawn from the Generalized Pareto distribution with cdf $F(x) = 1 - (1-x)^{c_B}$. Likewise, let the valuation of a given home to the seller be $\kappa + (1-\alpha) z_t + \alpha z_{t-1} + s$, where s is specific to the seller and is drawn independently from distribution $F(y) = 1 - (1-y)^{c_S}$, and κ is a constant. As before, $z_t = \mu + z_{t-1} + \eta_t$, with η_t white noise. Then the expected negotiated price is

$$p_t = z_t - \alpha\psi\eta_t + E_{b,s}\left[(1-\psi)\, b + \psi s \;\middle|\; b \geq s + \widetilde{\kappa} - \alpha\eta_t\right]$$

where the expectation is taken with respect to the joint distributions of b and s. The expectation is conditioned on the buyer valuing the property more than the seller. $\widetilde{\kappa} \equiv \kappa - \alpha\mu$ captures the degree of non-overlap in the supports of the distributions of buyers' and sellers' valuations.

Let $b^{(j)}$ indicate the jth highest value among the buyer-specific components of buyers' valuations. The expected price at auction, conditional on sale, is

$$a = z_t + \mathcal{H}(\eta_t) \tag{19}$$

where the function $\mathcal{H}$ varies according to how the auction price is modelled. In the first two scenarios, the seller chooses their reserve price non-strategically, setting it equal to their valuation. In the first scenario, the reserve price is not announced, and the bidders do not 'jump-bid' in order to exceed it; a transaction takes place then when the second highest bidder value exceeds the seller's valuation, and

$$\mathcal{H}(\eta_t) = E_{b,s:N}\left[b^{(2)} \;\middle|\; b^{(2)} \geq s + \widetilde{\kappa} - \alpha\eta_t\right]$$

In the second scenario, the seller reserve is announced, so that price is the maximum of the second highest buyer value and the reserve price, and a transaction takes place if the highest bidder's value exceeds the seller's valuation; thus

$$\mathcal{H}(\eta_t) = E_{b,s:N}\left[\max\left(b^{(2)}, s\right) \,\middle|\, b^{(1)} \geq s + \widetilde{\kappa} - \alpha\eta_t\right]$$

In the third scenario, the seller sets an optimal reserve price. Here, we need to take a stand on what sellers know about the buyer distribution. We assume that sellers know the shape of the buyer distribution but set the optimal reserve price as if the buyer common component were equal to their own, so that

$$\mathcal{H}(\eta_t) = E_{b,s:N}\left[\max\left(b^{(2)}, r(s)\right) \,\middle|\, b^{(1)} \geq s + \widetilde{\kappa} - \alpha\eta_t\right]$$

where $r(s) \equiv \frac{1+c_B(s+\widetilde{\kappa})}{1+c_B}$. First-differencing (19), and then taking a linear approximation of each price around a non-stochastic steady state, and ignoring constant terms, we have

$$\begin{aligned}
\Delta a_t &\approx \eta_t - \alpha q^a (\eta_t - \eta_{t-1}) \\
\Delta p_t &\approx \eta_t - \alpha (\psi + q^p)(\eta_t - \eta_{t-1})
\end{aligned}$$

where $q^a \equiv H'(0)$ and $q^p \equiv \frac{\partial E_{b,s}\left[(1-\psi)b+\psi s \middle| b \geq s+\widetilde{\kappa}+x\right]}{\partial x}_{|x=0}$. As before, the two series are cointegrated since a linear approximation of the difference between auction and private-treaty prices is stationary

$$a_t - p_t = \alpha\left[\psi - (q^a - q^p)\right]\eta_t$$

There is also a VECM representation, as before

$$\begin{aligned}
\Delta a_t &= \frac{q^a}{\psi - (q^a - q^p)}(a_{t-1} - p_{t-1}) + (1 - q^a\alpha)\eta_t \\
\Delta p_t &= \frac{\psi + q^p}{\psi - (q^a - q^p)}(a_{t-1} - p_{t-1}) + (1 - \alpha(\psi + q^p))\eta_t
\end{aligned}$$

Unlike the model in which all transactions are consummated, here each series Granger causes the other. Also, now both series display positive momentum, and

not only private-treaty prices

$$Cov\left(\Delta a_t, \Delta a_{t-1}\right) = \left(1 - q^a \alpha\right) q^a \alpha Var\left(\eta_t\right)$$
$$Cov\left(\Delta p_t, \Delta p_{t-1}\right) = \left(1 - \left(\psi + q^p\right)\alpha\right)\left(\psi + q^p\right)\alpha Var\left(\eta_t\right)$$

Unconsummated transactions thus rob our preferred model of its stark predictions for Granger causality and momentum. Quantitatively, however, those predictions can still hold. The relative behaviour of the two price indices depends on the ratios $r_1 \equiv \frac{q^a}{\psi+q^p-q^a}$ and $r_2 \equiv \frac{q^a}{\psi+q^p}$. If these are small, the combination of auctions weighting buyer valuations more heavily and a lagging seller valuation will continue to predict an auction price momentum substantial smaller than the private-treaties. It will also predict that the contribution of private-treaty prices to predicting auction prices is small relative to that of lagged auction prices to predicting private-treaty prices.

The ratios depend on the buyer and seller distributions, the number of bidders and ψ. Examining the two terms analytically is intractable, in large part due to the need to work with the second order statistic. However, simulations that assume equal bargaining power (consistent with our estimates in Table 11) and are calibrated by choosing a value of $\widetilde{\kappa}$ to fit the observed fraction of auctions that end in a sale, yield small ratios for a wide range of distribution shapes and for a sufficient number of bidders.

Figures 7 shows the ratios r_1 and r_2 for each scenario with three distributions cases: uniform $[c_B = c_S = 1]$, $[c_B = 2, c_S = 0.5]$ (right-skewed buyer, left-skewed seller) and $[c_B = 0.5, c_s = 2]$ (left-skewed buyer, right-skewed seller). We also report the empirical counterpart for r_2 based on $Cov\left(\Delta a_t, \Delta a_{t-1}\right)/Cov\left(\Delta p_t, \Delta a_{t-1}\right)$, as taken directly from the data.

Figure 7 highlights that with sufficiently many bidders (typically six is enough), all three scenarios can replicate the r_2 ratio observed in the data. That is, the autocorrelation in auction prices is small relative to their information content for private-treaty prices. Thus, even when allowing for unconsummated transactions, the results are still consistent with the main text's findings. This is also supported by the fact that r_1 is typically close to zero in the simulations, consistent with an insignificant coefficient on the adjustment parameter in the auctions equation, α_a,

from Table 5. Figure 8 further highlights that q^a is small, below 0.3 with six or more bidders, which guarantees the condition $\alpha q^a < 1$ (Section IV.E) for sellers only having a small overall effect on auction prices.

B. *Asymmetry in the Lead/Lag Relationship*

Table 12 shows correlations between auction price and private-treaty price growth, that condition on the previous direction of price changes (in levels or relative to the sample average). The top panel presents the correlation of past auction price growth with current private-treaty price growth, when past auction prices are rising in levels (or by more than their mean); the bottom panel presents the correlation of past private-treaty price growth with current auction price growth, when past private-treaty prices are falling (rising less than their mean). Column 1 reports the unconditional correlations using the entire sample.

From the first column, we see that, as expected, the top correlation is much greater than the bottom (auction prices Granger cause private-treaty prices but the reverse is not true). In the next two columns, we see that auction price growth does predict private-treaty price growth when auction prices are rising (the top panel), but that private-treaty price growth does not predict auction prices when private-treaty prices are falling or rising by less than their mean (the bottom panel). The absence of predictability using private-treaty prices, when previous price growth is negative (or rising less than usual), does not favour a model with asymmetry in the lead-lag relationship.

Table 12—Conditional Correlations: Auction and Private-treaty Prices

	$\rho(\Delta p_t, \Delta a_{t-1})$	$\rho(\Delta p_t, \Delta a_{t-1}) \mid \Delta a_{t-1} > 0$	$\rho(\Delta p_t, \Delta a_{t-1}) \mid \Delta a_{t-1} > \overline{\Delta a_{t-1}}$
Sydney	0.51	0.41	0.31
	(81)	(51)	(41)
Melbourne	0.59	0.63	0.58
	(77)	(54)	(36)
	$\rho(\Delta p_{t-1}, \Delta a_t)$	$\rho(\Delta p_{t-1}, \Delta a_t) \mid \Delta p_{t-1} < 0$	$\rho(\Delta p_{t-1}, \Delta a_t) \mid \Delta p_{t-1} < \overline{\Delta p_{t-1}}$
Sydney	0.11	-0.05	-0.10
	(81)	(24)	(39)
Melbourne	0.33	-0.39	0.13
	(77)	(19)	(44)

Note: Number of observations used in calculating correlation reported in parentheses.

As an alternative structural check, we also estimate non-linear models that explicitly allow for a kinked response in auction prices, conditioning on the direction of the change in the permanent component in price relative to its mean.[36] Retaining the assumption that the response in private-treaty prices to shocks is symmetric, as suggested by the theoretical discussion above, the non-linear model for the auction price is

$$\Delta a_t = \mu + \beta^+ \eta_t I\left(\eta_t \geq 0\right) + \beta^- \eta_t I\left(\eta_t < 0\right) + \varepsilon_t^a - \varepsilon_{t-1}^a$$

where $I\left(.\right)$ is a binary indicator function taking a value of one when the condition in its argument is satisfied and zero otherwise. The second model we consider allows both for asymmetry and lagged diffusion of shocks in the auction price

$$\begin{aligned}\Delta a_t = \mu &+ \left(1-\gamma^+\right)\eta_t I\left(\eta_t \geq 0\right) + \gamma^+ \eta_{t-1} I\left(\eta_{t-1} \geq 0\right) \\ &+ \left(1-\gamma^-\right)\eta_t I\left(\eta_t < 0\right) + \gamma^- \eta_{t-1} I\left(\eta_{t-1} < 0\right) + \varepsilon_t^a - \varepsilon_{t-1}^a\end{aligned}$$

For each model, estimation is undertaken jointly with private-treaty prices (Equation 13) and list prices (Equation 14), which assist in providing conditional identification of the unobserved permanent shock, η_t.

The results in Table 13 suggest there is little evidence in support of an asymmetric response in auctions prices to shocks in the permanent trend. Across all models, the contemporaneous point estimates for the response to a positive shock (β^+ or $1-\gamma^+$) are less than the response to a negative shock (β^- or $1-\gamma^-$). This is the opposite prediction to that implied by asymmetry and lagged diffusion of shocks across the buyer distribution, which predicts a stronger response to positive shocks than for negative. Accounting for the uncertainty around the point estimates, the null hypotheses of no asymmetry in the response to shocks ($\beta^+ = \beta^-$ and $\gamma^+ = \gamma^-$) cannot be rejected in either city. Overall, there is little evidence of lagged diffusion of shocks through the buyer distribution.

C. *Robustness of the Empirical Findings*

This section addresses the robustness of our core empirical findings to: (a) measurement of the hedonic indices; (b) selection of the sales mechanism; and (c) whether other housing characteristics, either observed or unobserved, could explain

[36] Similar results are obtained allowing for kinked effects around zero growth rather than mean growth.

Table 13—Asymmetry in the Response of Auction Prices to Shocks

	Contemporaneous asymmetry		Asymmetry and lagged diffusion	
	β^+	β^-	γ^+	γ^-
Sydney	0.73	1.58	0.48	-0.09
	(0.08,1.32)	(0.47,2.13)	(-0.08,0.91)	(-0.78,0.44)
Melbourne	1.07	1.12	0.56	0.42
	(0.67,1.60)	(0.62,1.58)	(-1.17,1.25)	(-0.72,1.31)
p-values	$H_0 : \beta^+ = \beta^-$		$H_0 : \gamma^+ = \gamma^-$	
Sydney	*0.17*		*0.27*	
Melbourne	*0.84*		*0.87*	

Note: Point estimates are computed using two-step maximum likelihood; bootstrapped percentile confidence intervals (at 95 per cent in parenthesis) and p-values (in italics) are reported; Melbourne data are adjusted for seasonality prior to estimation.

them. For brevity, we focus on in-sample results.

Measurement

To address whether measurement could be a concern, we revisit the in-sample Granger causality tests using data without attribute controls – specifically, the number of bedrooms, bathrooms and log size. We also revisit them using repeat-sales instead of hedonic indices, which effectively difference out unobserved time-invariant characteristics of homes.[37] Table 14 shows that our results are robust to the omission of attributes, and to using alternative repeat-sales indices. As such, our main findings do not appear to be sensitive to the measurement approach taken.

Table 14—In-Sample Causality Robustness: Varying Hedonic Controls and Repeat-Sales

Null hypothesis	Sydney (hedonic)	Sydney (repeat-sales)	Melbourne (hedonic)	Melbourne (repeat-sales)
$a_t \overset{GC}{\nrightarrow} p_t$	25.19***	15.62***	36.70***	24.16***
	(0.00)	(0.01)	(0.00)	(0.00)
$p_t \overset{GC}{\nrightarrow} a_t$	3.86	8.39	2.15	3.02
	(0.57)	(0.14)	(0.71)	(0.55)

Note: Using Toda and Yamamoto's (1995) testing approach. All tests include seasonal controls; the lag structure is unchanged from that used in the main text; p-values are reported in parentheses. For the hedonic indices only controls for the postcode and property type are included while the number of bedrooms, bathrooms and log size are omitted.

[37] It should be noted that a limitation of using repeat-sales is that they introduce scope for sample-selection bias since only multiple sales observations are used in their calculation. However, consistent with Hansen (2009), we find that the indices are comparable at the city-wide level, exhibiting similar price dynamics in terms of their leading and lagging properties, and in their average estimated growth rates.

Table 15 shows that for repeat-sales,the contribution of transaction level variance to the variance of quarterly growth is small.

Table 15—RMSE of Repeat-Sales Price Regressions

	RMSE	N	$\frac{RMSE}{\sqrt{N}}$	St. Dev.
		Sydney		
Auction	0.21	209	0.015	0.029
Private	0.23	6 166	0.003	0.017
		Melbourne		
Auction	0.22	334	0.012	0.033
Private	0.30	4 984	0.004	0.021

Note: RMSE is the root mean squared error of the corresponding regression; N is the average number of observations per quarter; and St. Dev. is the standard deviation of quarterly prices growth.

Selection of the sales mechanism

Another possible concern is that endogeneity in sellers' selection of sales mechanism could be responsible for the differential time series properties of auctions and private-treaty prices. To address this, we first provide reduced-form and then structural evidence that does not support a selection explanation.

If selection is driving our findings, then one might expect that selection of the sales mechanism could itself have predictive information for prices. For example, if sellers are forward looking and auctions are more profitable in rising markets then one might expect the share of auctions in total sales to pick up before prices rise. Conversely, the share of private-treaties will increase when prices fall.[38] To examine if this true, we augment the benchmark VAR with a measure of the auction share – the ratio of all auctions held (successful and unsuccessful) to all sales events held (i.e. successful and unsuccessful auctions plus private-treaty sales) and the auction clearance rate. We include the latter to separately control for predictability through the sales mechanism chosen and predictability through clearance (the fact that buyers' values update more quickly than sellers).[39]

Based on the evidence in Sydney (Table 16), the share of auctions held, as a proportion of all sales, is only useful for predicting the clearance rate and itself. It

[38] In addition, one might expect to the extent that both sellers and buyers are forward looking, and fully update their information in response to a common shock, the auction sales rate – the ratio of successful auctions to all auctions held – should not assist when predicting future price growth. This is true, for example, in Wang's (1993) dynamic model with endogenous selection.

[39] Without a control for the auction clearance rate, it is possible (likely) that auction incidence is correlated with the clearance rate and so omitting it could lead to spurious inference when buyers values update more quickly than sellers in response to shocks.

has no predictive content for future price formation as one might expect if selection is the explanation. Similar results are found in Melbourne where the auction share is also not informative for forecasting prices.

Table 16—Causality Tests with Auction Incidence and the Clearance Rate

H_0 : $\overset{GC}{\nrightarrow}$	Auction prices a_t	Clearance rate c_t	Private-treaty prices p_t	Auction incidence v_t
		Sydney		
a_t	0.48	0.00***	0.00***	0.00***
c_t	0.23	0.02**	0.00***	0.02**
p_t	0.53	0.02**	0.01***	0.03**
v_t	0.35	0.00***	0.37	0.00***
		Melbourne		
a_t	0.36	0.99	0.00***	0.35
c_t	0.03**	0.35	0.01**	0.04**
p_t	0.34	0.90	0.29	0.51
v_t	0.50	0.73	0.39	0.02**

Note: ***,** and * denote significance at the 1, 5 and 10 per cent levels; $H_0 : \overset{GC}{\nrightarrow}$ denotes the null of non-causality from the row variable to the column variable; p-values are reported.

Taking a structural approach, we also undertake Granger causality tests using estimated price indices that adjust for endogenous selection of the sales mechanism. Specifically, we use a two-step Heckman estimator of the endogenous switching regression

$$a_{ijt} = \begin{cases} 1 \text{ if } x'_{ijt}\kappa \geq u_{ijt} \\ 0 \text{ if } x'_{ijt}\kappa < u_{ijt} \end{cases}$$

$$\ln p_{ijt} = \begin{cases} \sum_{t=0}^{T} \gamma_t^a D_{it} + \sum_{j=1}^{J} \beta_j^a PC_{ij} + \sum_{k=1}^{k} \theta_k^a C_{ikt} + \varepsilon_{ijt}^a \text{ if } a_{ijt} = 1 \\ \sum_{t=0}^{T} \gamma_t^p D_{it} + \sum_{j=1}^{J} \beta_j^p PC_{ij} + \sum_{k=1}^{k} \theta_k^p C_{ikt} + \varepsilon_{ijt}^p \text{ if } a_{ijt} = 0 \end{cases}$$

where γ_t^a and γ_t^p are the estimated selection-adjusted price indices for auctions and private treaties at time t. Assuming the selection and pricing residuals are joint normal (and correlated), the selection vector x'_{ijt} includes all variables in the price equation and a dummy variable for whether the property was ever previously auctioned successfully. If it is true that the current price of a property is statistically independent of the previous mechanism used to affect its sale, this variable assists in providing conditional identification of auction incidence.

Table 17 reports the share of auctions and private-treaties that were previously sold at auction: about 62 (58) percent of auctioned properties were previously sold through auction but only 6 (9) per cent of private-treaties were previously sold at auction for Sydney (Melbourne). Table 18 reports the in-sample causality findings using the selection-adjusted indices and the benchmark price indices estimated on the same sample.[40] Although controlling for selection does reduce the value of the test statistic when the null is that auction prices do not Granger cause private-treaty prices, and increases the value of the test statistic when the null is that private-treaty prices do not Granger cause auction prices, there is still clear evidence to reject the null that auction prices do not Granger cause private-treaty prices, but little evidence to reject the null that private-treaty prices do not Granger cause auction prices.

Table 17—Correlation Between d^a_{ijt} and a_{ijt}

	Auction previously chosen	No auction previously chosen	Total
		Sydney	
Auction	37,679	23,181	60,860
	(0.62)	(0.38)	
Private-treaty	31,983	534,199	566,182
	(0.06)	(0.94)	
Total	69,662	557,380	
		Melbourne	
Auction	77,842	56,123	133,965
	(0.58)	(0.42)	
Private-treaty	40,258	411,035	451,293
	(0.09)	(0.91)	
Total	118,100	467,158	585,258

Note: Proportions are in parentheses and are with respect to the row total. For example, of all auctions held, 62 per cent of them were previously auctioned at some point in the sample (within the sample of repeat-sales).

Using Real Prices

Here we show that our results are robust to the use of real house price indices rather than nominal. Using city-specific CPI's to deflate nominal prices in each city, Table 19 compares the autocorrelation coefficients in nominal and real (deflated) prices growth. The autocorrelation coefficients are similar across the two measures and there is less momentum in auction prices growth than there is in private-treaty

[40] The definition of d^a_{ijt} requires us to restrict the sample to repeat sales only.

Table 18—In-Sample Causality: Accounting for Endogenous Selection

Null hypothesis	Sydney (2step-HESM)	Sydney (Hedonic)	Melbourne (2step-HESM)	Melbourne (Hedonic)
$a_t \overset{GC}{\nrightarrow} p_t$	35.78***	51.58***	51.64***	48.86***
	(0.00)	(0.00)	(0.00)	(0.00)
$p_t \overset{GC}{\nrightarrow} a_t$	11.02	5.65	3.32	8.93
	(0.14)	(0.58)	(0.51)	(0.26)

Note: Using Toda and Yamamoto's (1995) testing approach; 2step-HESM stands for a two step-estimator of the Heckman Endogenous Switching model; all models estimated on the repeat-sales sample.

prices growth. Table 20 re-examines our key causality and efficiency findings using real prices: the results are qualitatively unchanged.

Table 19—Momentum in Real and Nominal Prices

Autocorrelation coefficient	Sydney Nominal	Sydney Real	Melbourne Nominal	Melbourne Real
		Auction price growth		
Lag 1	-0.05	-0.01	0.44	0.41
Lag 2	0.09	0.10	0.25	0.26
Lag 3	0.10	0.10	0.00	-0.02
Lag 4	0.01	-0.04	-0.14	-0.16
		Private-treaty price growth		
Lag 1	0.13	0.12	0.40	0.31
Lag 2	0.22	0.21	0.43	0.39
Lag 3	-0.03	-0.04	0.05	0.00
Lag 4	-0.08	-0.12	0.07	0.05

Note: Autocorrelations based on nominal data are the same as those reported in Figure 2 and are reported using all attribute controls for Sydney on the sample 1992:I to 2012:IV and limited attribute controls for Melbourne on the sample 1993:I to 2012:IV. City-specific CPI's for Sydney and Melbourne are sourced from the Australian Bureau of Statistics, Catalogue 6401.0, Table 5.

Housing characteristics and location

Our final checks examine whether the causality findings could be explained by the location or types of homes sold, rather than the mechanism used to sell it. Auction prices might capture more timely information because they are more prevalent in locations that lead housing prices. For example, inner city prices could lead middle and outer city prices and auctions are more frequent in inner city areas.

To address this concern, we estimate 14 (10) pairs of sub-city hedonic indices for Sydney (Melbourne), one index for auctions and one for private-treaties in each sub-city district.[41] Using a panel-VAR framework, which allows for heterogeneous

[41] We use Statistical Area Level 4 localities as defined by the Australian Bureau of Statistics. They are

Table 20—In-sample Granger Causality Tests: Real Prices

	All controls	
Null hypothesis	Sydney	Melbourne
$a_t \overset{GC}{\nrightarrow} p_t$	68.89***	11.95**
	(0.00)	(0.02)
$p_t \overset{GC}{\nrightarrow} a_t$	8.67	4.33
	(0.12)	(0.36)
$E\left(a_t \mid \mathcal{I}_{t-1}^{a,p}\right) = a_{t-1}$	11.67	9.97
	(0.23)	(0.19)
$E\left(p_t \mid \mathcal{I}_{t-1}^{a,p}\right) = p_{t-1}$	78.84***	28.18***
	(0.00)	(0.00)

Note: ***, ** and * denote significance at the 1, 5 and 10 per cent levels. $\overset{GC}{\nrightarrow}$ is a test for non-Granger causality; a_t is the real auction price, p_t the real private-treaty price and $\mathcal{I}_{t-1}^{a,p}$ is the information set at time $t-1$ (conditioning on lagged real auction and real private-treaty prices). All controls includes the property type and interactions with the number of bedrooms, the number of bathrooms, and the logarithm of the size of the property. For Melbourne this sample is restricted to 1997:IV onwards and includes controls for seasonality; p-values are in parentheses.

lagged diffusion and contemporaneous correlation in shocks across districts, we test whether the causality from auctions to private-treaty prices holds at this finer level of geographic disaggregation.[42] If location is the alternative explanation, we should not expect to find the same results to hold *within* sub-city districts.

Table 21 reports the results of panel-VAR Granger causality tests (Dumitrescu and Hurlin (2012)). They highlight that the null of non-causality from auction to private-treaty prices, within districts, is clearly rejected for both cities. There is little evidence to suggest that private-treaty prices similarly Granger cause auction prices, with the exception of perhaps Melbourne where there is a marginally significant p-value (0.09). However, this result is driven by several outer-city districts with very low district-specific auction shares (in the order of 1 to 2 per cent of all

areas with common socio-economic demographics and have 1200 (1350) private-treaty sales and 120 (220) auctions per quarter in Sydney (Melbourne) on average.

[42] The model panel-VAR is

$$\ln a_{j,t} = \sum_{k=1}^{K} \phi_{j,k}^{a} \ln a_{j,t-k} + \sum_{k=1}^{K} \phi_{j,k}^{p} \ln p_{j,t-k} + \varepsilon_{j,t}^{a}$$
$$\ln p_{j,t} = \sum_{k=1}^{K} \lambda_{j,k}^{a} \ln a_{j,t-k} + \sum_{k=1}^{K} \lambda_{j,k}^{p} \ln p_{j,t-k} + \varepsilon_{j,t}^{p}$$

for all sub-regions $j = 1, ..., J$ with $E\left(\varepsilon_t \varepsilon_\tau'\right) = \Sigma_\varepsilon$ for $t = \tau$ and $\mathbf{0}$ for $t \neq \tau$ $\left(\varepsilon_t \equiv \left[\varepsilon_t^a, \varepsilon_{,t}^p\right]', \varepsilon_t^a \equiv \left[\varepsilon_{1,t}^a ..., \varepsilon_{J,t}^a\right], \varepsilon_{,t}^p = \left[\varepsilon_{1,t}^p, ..., \varepsilon_{J,t}^p\right]\right)$. The null hypotheses are $p_t \overset{GC}{\nrightarrow} a_t$ ($H_0 : \phi_{j,k}^p = 0$ for all $j = 1, ..., J$ and $k = 1, .., K-1$) and $a_t \overset{GC}{\nrightarrow} p_t$ ($H_0 : \lambda_{j,k}^a = 0$ for all $j = 1, ..., J$ and $k = 1, ..., K-1$) with at least one non-zero element under the alternative in each case.

sales within a district), and that comprise a low share of total sales overall. These districts are heavily affected by small-sample noise in the underlying auction price estimates, which obscures the underlying relationship in price across mechanisms.[43] Restricting attention to districts where auctions comprise a minimum of 7.5 per cent of all sales (within the district) – thus mitigating the small sample concern – there is strong evidence to suggest auction prices Granger cause private-treaty prices, but no evidence to suggest that private-treaty prices are similarly informative.

Table 21—In-Sample Causality Conditioning on Sub-city Prices

Null hypothesis	Sydney		Melbourne	
	All districts	Min. auction share	All districts	Min. auction share
$a_t \overset{GC}{\nrightarrow} p_t$	4.27***	6.16***	4.14**	4.56***
	(0.01)	(0.00)	(0.02)	(0.01)
$p_t \overset{GC}{\nrightarrow} a_t$	8.71	4.95	2.81*	2.03
	(0.21)	(0.32)	(0.09)	(0.20)
No. of districts	14	6	10	6

Note: Test statistics are the $Z_{N,T}^{Hnc}$ test-statistic with residual bootstrapped p-values in parentheses to account for cross-locality error dependence (Dumitrescu and Hurlin (2012)). All districts denotes Statistical Area Level 4 localities in Sydney and Melbourne as defined by the Australian Bureau of Statistics and are based on the 2011 concordance (1270055006C183 Postcode to Statistical Area Level 4). Min. auction share restricts attention to districts where at least 7.5 per cent of successful sales are auctions.

Finally, conditioning on housing type, Table 22 revisits our in-sample causality findings using houses in lieu of auctions, and units (apartments) in lieu of private-treat sales. If houses are a better guide as to future market conditions, we would expect them to Granger cause unit prices, but that the reverse would not be true. The data are inconsistent with this hypothesis: in both cities there is bivariate causality between house and unit prices. In contrast, if we focus on within-group variation (either house sales or apartment sales), we still find unidirectional causality from auction to private-treaty prices by the type of home sold (Table 23).

VII. Cointegration Results

Table 24 reports results from bivariate and trivariate cointegration tests using Johansen's Likelihood Ratio (Trace) Test. At conventional levels of significance, all tests are consistent with the presence of a single common trend in price.

[43] This is also reflected in graphs of prices where auction prices still lead private-treaty prices, but the former's volatility obscures the presence of a leading relationship when testing for Granger causality. The same effect is also present in Sydney and is reflected in a thicker right tail of the test statistic distribution under the null that private-treaty prices Granger cause auction prices, than under the null that auction prices Granger cause private-treaty prices.

Table 22—In-Sample Causality: Do House Prices Lead Apartment Prices?

Null hypothesis	Sydney All-sales	Melbourne All-sales
$h_t \overset{GC}{\nrightarrow} u_t$	57.27***	18.16***
	(0.00)	(0.00)
$u_t \overset{GC}{\nrightarrow} h_t$	9.18*	18.10***
	(0.10)	(0.00)

Note: h_t denotes house prices and u_t apartment prices.

Table 23—In-Sample Causality Conditioning on The Type of Housing Sold

Null hypothesis	Sydney houses	Sydney apartments	Melbourne houses	Melbourne apartments
$a_t \overset{GC}{\nrightarrow} p_t$	60.20***	5.75**	34.38***	7.27**
	(0.00)	(0.02)	(0.00)	(0.03)
$p_t \overset{GC}{\nrightarrow} a_t$	5.31	0.48	1.60	3.87
	(0.27)	(0.49)	(0.81)	(0.14)

Note: Tests based on the apartments sample are restricted to 1997 onwards due to the small number of auctioned apartments prior to that date.

VIII. Copyright and Disclaimer Notices

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Table 24—Johansen Trace Test Results

Variables	No cointegration	Null Hypothesis 1 cointegrating vector	2 cointegrating vectors
		Sydney	
s_t and a_t	20.20***	3.03	–
s_t and p_t	28.27***	3.14	–
a_t and p_t	21.39***	3.29	–
a_t, p_t and s_t	50.96***	20.62***	3.42
		Melbourne	
s_t and a_t	18.58**	0.92	–
s_t and p_t	23.70***	0.45	–
a_t and p_t	20.44***	0.89	–
a_t, p_t and s_t	63.70***	18.27**	2.28

Note: *** and ** denote rejection of the null at the 1 and 5 per cent levels of significance. Tests are in-sample and based on Johansen's Trace Test statistic.

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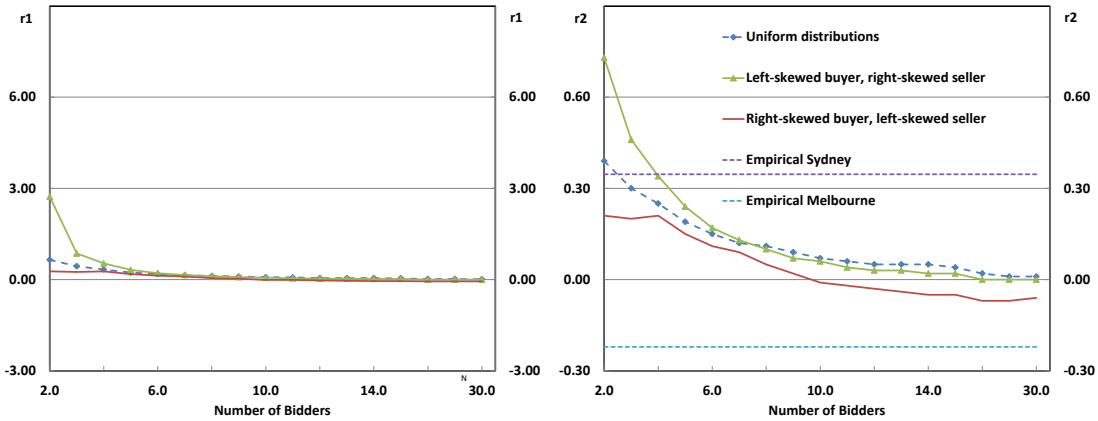
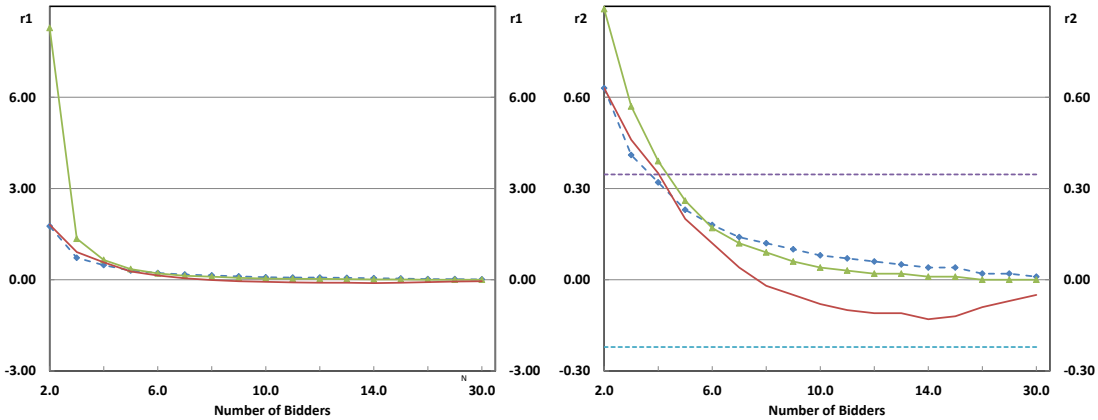
Scenario 1: Hidden Reserve Price



Scenario 2: Announced Reserve Price



Scenario 3: Optimal Announced Reserve Price

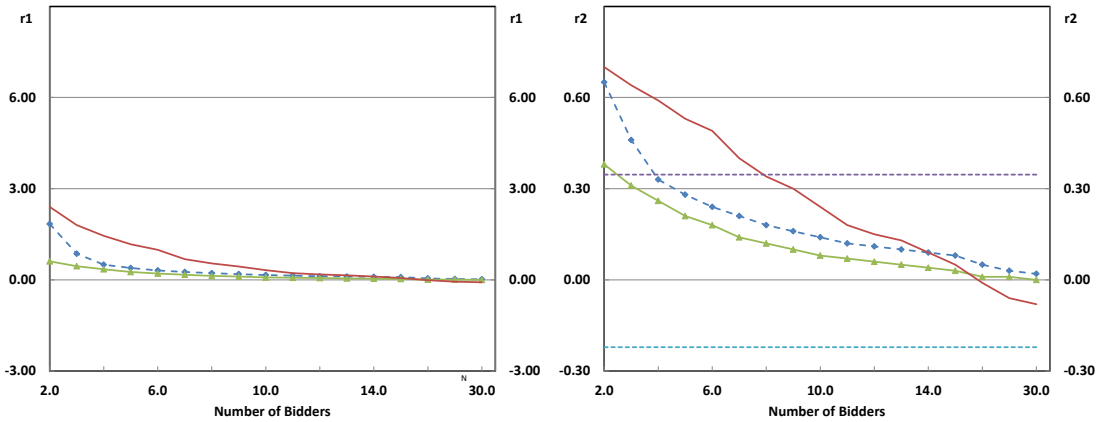
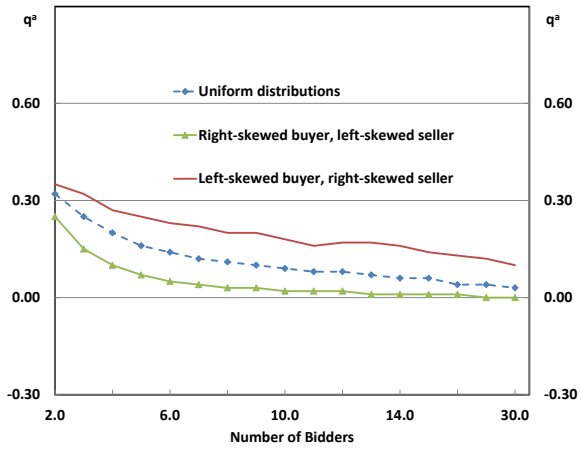


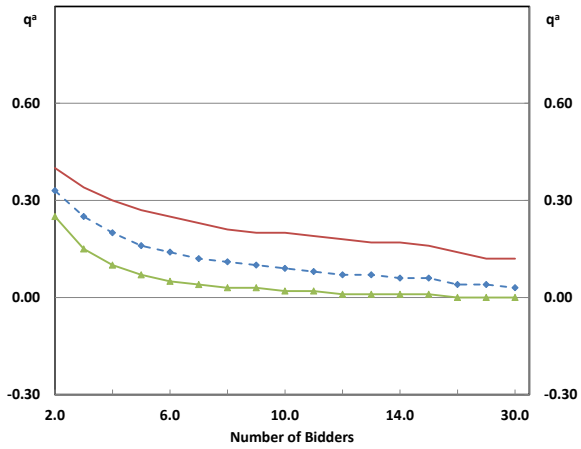
FIGURE 7. COMPARISON OF SIMULATED AND EMPIRICAL PRICE RATIOS

Note: The empirical "r2" for Sydney and Melbourne are calculated using the ratio of $cov\left(\Delta a_t, \Delta a_{t-1}\right)/cov\left(\Delta p_t, \Delta a_{t-1}\right)$ based on a comparable sample from 1997:II to 2012:IV using the hedonic indices with all attributes (and their interactions with the property type) estimated.

Scenario 1: Hidden Reserve Price



Scenario 2: Announced Reserve Price



Scenario 3: Optimal Announced Reserve Price

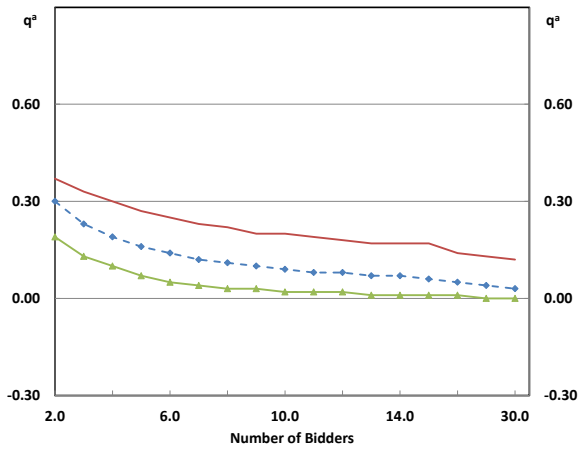


FIGURE 8. THE MAGNITUDE OF q^a