The Unilateral Price Effects of Several National SFR Mergers

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Abstract

Reporting around the recent rise in mergers among large players in the single family rentals (SFR) space has mostly emphasized the increase in market power that these merging firms would be able to exploit to raise rents. On the other hand, public communications around these mergers emphasize the efficiency gains that would be achieved through spatial density of ownership. Using a 2-way fixed effects regression, I estimate the price effects of 3 large SFR mergers, and use the results from a simple theoretical model to draw implications for the merger-induced changes in cost efficiency. I show that the low levels of concentration and pre-merger market shares result in very modest and nearly economically insignificant changes in post-merger rent prices for most local markets, with estimated price increases and decreases in even the most concentrated markets of under 10% and 20%, respectively. I also show that there is substantial heterogeneity in net price effect both across mergers and across pre-merger firm ownership – namely whether a property is owned by the merger target, acquiror, or outside firm results in price changes of different magnitude and sign. My results demonstrate the importance of considering heterogeneous price responses and the influence of merger-induced efficiencies when conducting ex-post merger evaluation.

1 Introduction

Following the financial crisis of 2008, an unprecedented quantity of distressed homes entered the housing market, which were in turn quickly bought up by Real-Estate Investment Trusts (REITs), bespoke entities formed by private equity firms with the primary goal of investing in these discounted homes. These trusts took these heavily foreclosed homes and refashioned them as rentals, funded through debt and equity securities backed by the generated rental income (Abood, 2017). In recent years, REITs have also turned to M&A as an alternative method of portfolio growth and capturing returns to scale in order to meet the financial demands of their shareholders.

This "securitization of suburbia" following the foreclosure crisis of 2008-2010 is a phenomenon that has been commented on by news outlets for quite some time[1], yet the concern that these investment trusts are exploiting their amassed local market power to raise rents in order to offer better returns to investors has not been well-studied empirically. Despite media scrutiny and public outcry over the poor business practices of some of these institutional investors, little work has been done to establish a systematic relationship between consolidation and rent price in SFRs. Furthermore, as has been illustrated extensively in the mergers literature, it is ex-ante unclear what the net welfare effects of consolidation are. Though in general mergers between competing firms in differentiated products industries generate a gross upwards pricing pressure (GUPPI), this upwards pressure on price can be counterbalanced by the competitive response and marginal cost efficiencies resulting from the merger (Farrell and Shapiro, 2010). Larger institutional owners may also play a role in increased neighborhood gentrification, resulting in better amenities, infrastructure, and living standards. Wu et al. find that while mergers among institutional SFR investors result in increased rents, there is also a positive neighborhood effect of more concentrated institutional ownership in terms of reduced crime (Wu et al., 2020).

In this paper, I attempt to elucidate the relationship between consolidation of institutional ownership in local SFR markets and rent price. I first motivate some basic intuition by presenting a simple 3-firm model of horizontal mergers in a rental market setting, and solve for the post-merger equilibrium price change under general linear demand and cost efficiencies. The model results in certain implications around post-merger marginal cost reductions depending on the direction of the realized price effects.

At the heart of this study are my empirical estimates of the unilateral price effects of each SFR merger, using a 2-way fixed effects specification that exploits the variation in pre-merger local market structures, within-market firm ownership, and prices to identify heterogeneous price effects from merger-induced increases in market concentration for the merger targets, acquirors, and other market participants. I find significant but very modest price effects resulting from these mergers, with heterogeneity in the magnitude and sign of the effects depending on pre-merger firm ownership. This heterogeneity is suggestive of differences in merger-induced cost efficiencies. Applying these estimates to the results from my theoretical

[1] https://www.mpamag.com/news/marketing/reotorentals-wall-street-meets-main-street-13058.aspx
https://www.nytimes.com/2016/06/27/business/dealbook/private-equity-housing-missteps.html

model, I am able to make certain weak predictions regarding the relative magnitude of marginal cost efficiencies for each merging firm that would be necessary in order to observe the price changes I measure. I conclude by discussing the assumptions necessary for my conclusions to hold, and outline future steps for extending this paper, such as structural estimation and welfare analysis, in order to address the still-outstanding question of the net welfare effects of these mergers.

This paper contributes to a number of different streams of literature. First, my findings add to the rapidly growing body of work documenting the growth of the institutional SFR market and its consequences for investors and consumers (Wu et al., 2020; D'Lima, 2019; Ganduri et al., 2019). In particular, this paper can be seen as a complement to a working paper by Wu et al, who also use a 2-way FE specification to estimate the price effects, among other outcomes, of some of the same mergers among institutional SFR investors that I cover in this study[2]. The measure of rent Wu et al use is Zillow's monthly rent index, and they pool all mergers into a single DiD equation, such that the interpretation of their treatment coefficient is the average effect of a merger on the median rent in a treated neighborhood. Wu et al. find that the SFR mergers result in modest neighborhood price increases, though also result in decreases in neighborhood crime rate. This study aims to improve on the precision and expand the economic interpretation of Wu et al's inital price estimates by exploiting a richer dataset. If the market power mechanism of post-merger price change is true, then we ought to be able to estimate a measure of price change per unit of merger-induced concentration. This measure would be of greater economic interest than a general lump-sum price effect of a merger, which is what Wu et al. estimate.

Wu et al. also estimate a version of their main regression with an interaction with post-merger concentration HHI. As emphasized by Dafny et al., however, there are plausible endogeneity concerns with post-merger HHI and prices (Dafny et al., 2012). These concerns are especially plausible in the rental market context, in which local economic conditions can drive substantial change in the general makeup and preferences of the renting population. I address this in my own estimates by instead using pre-merger market shares as a reduced-form instrument for merger-induced change in concentration. Though I am able to reproduce roughly the same magnitude of price effects as Wu et al. in my pooled specification, I also show that selection into significant pre-merger trends biases the initial estimates, and there is important heterogeneity in effect direction both across mergers as well as particular ownership of merge-affected properties. Finally, Wu et al. use a limited pre and post period of 12 months for their estimates, so it is likely that they only capture the short run price outcomes of these mergers.

In short, the granularity and coverage of my data will allow me to provide more precise and economically meaningful estimates, and thus draw a more nuanced conclusion regarding the effect of SFR mergers on prices. I show through my results that the market power story

[2]When formulating my idea for this paper, the Wu et al. paper was a largely incomplete working draft, with a different estimation framework. It was only until I was nearly done writing this paper that Wu et al. published an updated version of their paper with a more similar specification to the one I propose in this paper. Regardless, I believe there are key differences in the framework I introduce in this paper that offer important new insights that Wu et al do not address.

around price change needs to be considered alongside the possible marginal cost efficiencies and particular pre-merger competitive structure in order to grasp the overall picture surrounding SFR mergers.

This paper also contributes more generally to the large body of extant literature aimed at estimating the ex-post price and welfare effects of horizontal mergers (Miller and Weinberg, 2017; Ashenfelter et al., 2015). To my knowledge, this paper along with the paper by Wu et al. are the only existing attempts to estimate these effects in the context of the institutional SFR market. Unfortunately without a formal demand estimation, whether these mergers ultimately benefited or harmed consumers is still an open question. Regardless, this paper takes an important step in adding nuance to the conversation around market power and prices by pointing out that cost efficiencies likely play a significant role in mediating the effects of an increase in concentration on post-merger prices.

2 Institutional Background

Though institutional investment in single-family homes started as early as 2011, the federal government played a major role in spurring its growth in 2012 through the creation of the original REO-to-Rental pilot program spearheaded by the Federal Housing Finance Agency (FHFA). The goal of this program was to revitalize the housing market by stimulating investor appetite for the large inventories of foreclosed homes, known as real-estate owned (REO), that resulted from the financial crisis. The program would "allow qualified investors to purchase pools of foreclosed properties with the requirement to rent the purchased properties for a specified number of years" [3]. Other government leaders also stepped in to encourage private investment into the single-family rental space, such as then-chair of the Federal Reserve Ben Bernanke, who stated that "rental housing markets having strengthened somewhat...some REO holders might come out ahead by renting, rather than selling, some of their properties" in a speech at the 2012 National Association of Homebuilders International [4]. Since 2011, the number of renting households has increased from 38 million to 44 million, while investors have spent over $40 billion on over 240,000 homes[5].

Despite this rapid growth in investment, institutional investors still only make up a small share of the total market for rental homes, around 2% in 2017[6]. That hasn't prevented the flurry of media reporting of the experience of disgruntled renters under these institutional investors. According to several articles spanning the Atlantic to the New York Times, these investors have engaged in increasingly draconian cost-cutting measures, such as passing on more maintenance responsibilities onto the tenant in 40+ page-long leases, aggressively evicting tenants, and tacking on various charges and fines to bills, on top of raising rents[7].

[3] https://www.fhfa.gov/Media/PublicAffairs/Pages/FHFA-Announces-Interested-Investors-May-PreQualify-for-REO-Initiative.aspx

[4] https://www.federalreserve.gov/newsevents/speech/bernanke20110210a.htm

[5] https://www.amherstcapital.com/documents/20649/0/NREI+-+Single-Family+Rentals+-+JAN+2019/2a39ed33-4cfd-411c-99bb-d60ebe0299b2

[6] https://www.urban.org/urban-wire/five-things-might-surprise-you-about-fastest-growing-segment-housing-market

[7] `https://www.nytimes.com/2020/03/04/magazine/wall-street-landlords.html`

Much of the public discourse at least partially attributes these reported abuses to the local market power that large institutional investors have accrued through a recent pivot towards mergers that have been driving up firms' neighborhood presence. Indeed, public communications from these firms emphasize the "operating benefits of local density as well as economies of scale" of these acquisitions[8]. However, there is ex-ante nothing to suggest that these claims to efficiency are not true. The primary recurring costs that rental companies face are maintenance costs, tenant services, property taxes, and insurance, of which only maintenance costs and tenant services are adjustable (Goodman, 2004). It is easy to see how larger rental companies could make use of geographic density and centralized property management platforms to reduce these costs.

I aim to address in this study a narrow segment of the issues raised above – specifically the question of whether the recent mergers among institutional SFR owners have caused increased rents through the market power channel, and consequently whether the claims to increased operational efficiency can be supported by the data.

3 A model of horizontal mergers in rental markets

In this section, I present a static model of a 3-firm Bertrand oligopoly in order to inform intuition for my later empirical estimates. Though simple, this model will allow me to take into consideration differences in rental home characteristics and cost efficiencies across firms, as well as characterize post-merger price changes as a function of pre-merger prices, quantities, and merger-induced cost efficiencies. This will allow me to back out implications regarding the likely direction and magnitude of marginal cost changes using the price effects I estimate in my empirical analysis.

To formalize, we have 3 rental companies which each supply a differentiated "housing service", in this case a single family home for rent. The demand for a home from company i is given by $D_i(p_i, \boldsymbol{p}_{-i})$, where $\boldsymbol{p}_{-i}$ is the vector of prices of the other two types of homes in the market. We will assume constant marginal costs, with each firm's cost function given by $c_i(D_i(p_i, \boldsymbol{p}_{-i}))$. The constant marginal costs assumption is one commonly assumed in the empirical literature, and I argue may reasonably be applied in the rental housing setting, since as mentioned in Section 2, the primary operating costs facing rental home owners are property taxes, insurance, maintenance, and tenant services, all of which should be fairly constant for each additional property of the same type. Profit from home type i will be

$$\pi_i = p_i D_i(p_i, \boldsymbol{p}_{-i}) - c_i(D_i(p_i, \boldsymbol{p}_{-i}) \tag{1}$$

The first order conditions that define a Nash equilibrium in this setting are

$$\frac{\partial \pi_i}{\partial p_i} = D_i + (p_i - c_i') \frac{\partial D_i}{\partial p_i} = 0 \tag{2}$$

https://www.theatlantic.com/technology/archive/2019/02/single-family-landlords-wall-street/582394/
https://archive.curbed.com/2018/5/18/17319570/wall-street-home-rentals-single-family-homes-invitation

[8] https://www.sec.gov/Archives/edgar/data/1687229/000119312517029042/d260125d424b4.htm

where I drop the function notation for simplicity. Because I assume constant marginal costs, c_i' is the same regardless of the supply of houses. These first-order conditions can be rewritten to provide a well-known necessary condition for equilibrium

$$m_i = \frac{-1}{\epsilon_{ii}} \tag{3}$$

where $m_i = \frac{p_i - c_i'}{p_i}$ is the price-cost margin for house i, and $\epsilon_{ij} = \frac{\partial D_i / D_i}{\partial p_j / p_j}$ is the elasticity of substitution.

Without loss of generality, suppose firms 1 and 2 merge. The new post-merger first-order derivatives are given by the following

$$\frac{\partial \pi_1}{\partial p_1} = D_1 + (p_1 - c_1'^*)\frac{\partial D_1}{\partial p_1} + (p_2 - c_2'^*)\frac{\partial D_2}{\partial p_1} \tag{4}$$

$$\frac{\partial \pi_2}{\partial p_2} = D_2 + (p_2 - c_2'^*)\frac{D_2}{\partial p_2} + (p_1 - c_1'^*)\frac{\partial D_1}{\partial p_2} \tag{5}$$

$$\frac{\partial \pi_3}{\partial p_3} = D_3 + (p_3 - c_3')\frac{D_3}{\partial p_3} \tag{6}$$

where $c_1'^*$, $c_2'^*$ represent the post-merger marginal costs for home types 1 and 2, respectively.

Froeb et al. demonstrate that though Newton's method can be applied to approximate a generic analytical solution to the post-merger FOCs, evaluation of the derived solution depends on the second-order derivatives of the demand functions, i.e. the curvature properties of demand (Froeb et al., 2005).

Though getting precise post-merger equilibrium price effects requires assumptions on the concavity of the demand functions, past literature has shown that computing the marginal cost reductions necessary to keep prices constant depend on only first derivatives, and pre-merger prices and quantities (Werden, 1996; Stennek and Verboven, 2001). These "compensating marginal cost reductions" (CMCRs) are given by

$$\Delta \boldsymbol{c}_{comp} = -\boldsymbol{B}^{-1}\boldsymbol{z}(\boldsymbol{p}^0)|_{\boldsymbol{c}_{pre}} \tag{7}$$

where $\boldsymbol{B} = b_{ij}$, where $b_{ij} = -\frac{\partial D_j}{\partial p_i}$ if firm i owns house type j following the merger and 0 otherwise. The first order derivatives are evaluated at the pre-merger marginal costs $\boldsymbol{c}_{pre}$. In the 3-firm case with a merger between firms 1 and 2, this expression evaluates to

$$\Delta \boldsymbol{c}_{comp} = \frac{1}{\epsilon_{22}\epsilon_{11} - \epsilon_{21}\epsilon_{12}} \begin{pmatrix} \epsilon_{21}(\frac{D_2}{D_1} m_2 p_2 \epsilon_{22} - m_1 p_1 \epsilon_{12}) \\ \epsilon_{12}(\frac{D_1}{D_2} m_1 p_1 \epsilon_{11} - m_2 p_1 \epsilon_{21}) \\ 0 \end{pmatrix} \tag{8}$$

where m_i and ϵ_{ij} are the price-cost margin and price elasticity as previously defined. The details for the derivation of this result can be found in Appendix C. Observe that the compensating marginal cost reductions are increasing in magnitude with closeness of substitution between the two products, pre-merger prices, and pre-merger margins, and that the characteristics of non-merging firms do not enter at all into this formulation.

Though informative in terms of setting a soft lower bound on the marginal cost efficiencies a merging firm would need to experience on each of its products to result in a price reduction, (8) does not offer any insight for thinking about possible differences in price effects experienced by the merger participants and outside firms. To get at these differential effects, I follow Hausman et al. and assume a general linear form of demand, and solve the system of equations given by the post-merger FOCs (Hausman et al., 2011). Unlike Hausman et al., I allow for post-merger change in marginal cost and price of the non-merging housing type to adjust as well, and thus derive the full general equilibrium price effects under the 3-firm case. The derived price effects are:

$$\frac{\Delta p_1}{p_1} = \frac{1}{p_1 A}\Bigg[(\Delta c_1' + m_2^* p_2 d_{12})(4 - d_{23}d_{32}) + (\Delta c_2' + m_1^* p_1 d_{21})\left(2d_{12} + \frac{m_1 p_1 Q_2}{m_2 p_2 Q_1}(2d_{21} + d_{23}d_{31})\right)\Bigg] \tag{9}$$

$$\frac{\Delta p_2}{p_2} = \frac{1}{p_2 A}\Bigg[(\Delta c_2' + m_1^* p_1 d_{21})(4 - d_{13}d_{31}) + (\Delta c_1' + m_2^* p_2 d_{12})\left(2d_{21} + \frac{m_2 p_2 Q_1}{m_1 p_1 Q_2}(2d_{12} + d_{13}d_{32})\right)\Bigg] \tag{10}$$

$$\frac{\Delta p_3}{p_3} = \frac{m_3}{Q_3 A}\Bigg[\frac{1}{p_2}(\Delta c_2' + m_1^* p_1 d_{21})\left(\frac{2Q_2}{m_2}d_{23} + \frac{Q_1}{m_1}\left(\frac{p_2}{p_1}d_{12} + \frac{m_1 Q_2}{m_2 Q_1}d_{21}\right)\right) + \frac{1}{p_1}(\Delta c_1' + m_2^* p_2 d_{12})\left(\frac{Q_2 p_1}{m_2 p_2}d_{21}d_{23} + \frac{Q_1}{m_1}(2d_{13} + d_{12}d_{23})\right)\Bigg] \tag{11}$$

where

$$A = 8 - d_{21}\frac{m_1 p_1 Q_2}{m_2 p_2 Q_1}(2d_{21} + d_{31}d_{23}) - d_{12}\frac{m_2 p_2 Q_1}{m_1 p_1 Q_2}(2d_{12} + d_{13}d_{32}) - d_{13}(\frac{2m_3}{Q_3}d_{31} + d_{32}d_{21}) - 4d_{21}d_{12} - 2d_{23}d_{32} - d_{23}d_{12}$$

$d_{ij} = -\frac{\partial Q_j/\partial p_i}{\partial Q_i/\partial p_i}$ is the diversion ratio from house type i to house type j when the price of house type i increases. $\Delta c_i' = c_i'^* - c_i'$ is the post-merger change in marginal cost for house type i. $m_i^* = \frac{p_i - c_i'^*}{p_i}$ is the price-cost margin for house type i at pre-merger prices and post-merger marginal costs. The solution procedure for the 3-firm case is analogous to that presented for the 2-firm case in Hausman et al (Hausman et al., 2011).

It is clear from the above derivation how differences in merger-induced cost efficiencies, substitutability between house types, and pre-merger prices can drive heterogeneity in post-merger price effects. Though the solutions in (9)-(11) are likely not particularly reliable for getting precise estimates, there are some interesting implications for marginal cost that can be derived under different assumptions on the signs of the effects. For instance, if the observed % change in the price of house type 1 is 0 and the % change in the price of house type 2 is negative, i.e. $\frac{\Delta p_1}{p_1} = 0$, $\frac{\Delta p_2}{p_2} < 0$, then without any other assumptions we get the following inequalities

$$\Delta c_2' < -d_{21} m_1^* p_1 \tag{12}$$

$$\frac{\Delta p_3}{p_3} < 0 \tag{13}$$

This implies very strong cost efficiencies for house type 2 are necessary in order to observe no change in the price of house type 1 and a decrease in price of house type 2 following a merger. For instance in the 3-firm case with symmetric diversion ratios, i.e. $d_{ij} = 0.5\ \forall i, j$, the inequality on $\Delta c_2'$ implies that the merged firm experiences a marginal cost reduction on house type 2 greater than 50% of the post-merger per-unit profit on house type 1. Furthermore, in the above case the price of the outside house type will always decrease, though the relative magnitude will depend on the specific pre-merger conditions. The inequalities are reversed if instead the price of house type 2 increases.

I will show in my empirical estimates that the price effects of the SFR mergers demonstrate the heterogeneity among market participants postulated by this simple model. This will allow me to make certain predictions regarding the mechanism and magnitude of the cost efficiencies realized by each merger.

4 Data and Sample Construction

4.1 Rent and Property Data

This paper makes use of rent data from 2 different sources, Zillow's proprietary Zillow Observed Rent Index (ZORI) and the Multiple Listing Service (MLS) dataset from CoreLogic. ZORI is a smoothed measure of the monthly median market rent. As of the time of this study, Zillow publishes ZORI estimates for the top 100 largest metropolitan areas in the United States, from January 2014 onwards, on varying levels of geography. In this study I use ZORI data published on the most granular geography at the time of collection, which are zip codes. By construction, ZORI values are directly comparable to monthly rents.

The MLS rent data is secured from CoreLogic and provides property-level list and close prices from its member MLS organizations. Data is available for all 50 states and each property is identified through a unique APN id. The data roughly spans the years 2007-2018, and annual rent is aggregated for each property by taking means.

Individual property ownership and characteristics data is taken from CoreLogic's tax assessor and deeds data. The tax assessor data is derived from tax rolls supplied by county governments, and the deed data is collected from various county clerk offices. The tax data covers the years 2007-2018 and the deeds data spans 30+ years up until 2019, and data on all 50 states is available. The deeds dataset is deduplicated to annual property observations and merged with the annual tax dataset prior to analysis. The combined dataset is then imputed over a fixed year range from 2000-2019, so that the final property dataset is a balanced panel of property-years. The details for this imputation procedure can be found in

Appendix D.

The property dataset is then merged with the MLS data by APN and year. Most of the properties in the sample only have reported rent prices for 1 or 2 years. This is not surprising, given that the MLS data consists of flow data from when a property is actually listed for sale by an MLS organization, so it is unlikely to observe prices for every year for every property. I assume in my empirical analysis that the linked MLS data consists of representative cross-sectional samples of prices for every market we observe in the sample.

4.2 Identifying Institutional SFR Investors

I make use of both web search and the Thomson ONE database to find M&A activity among large institutional SFR investors between 2014 and 2017. I find 5 mergers of interest, and summarize them in Table 1. I combine company filings from the SEC with the list of subsidiaries compiled by Abood '17 for private companies, and match subsidiary names against owner names in the property data, using both direct string matching as well as a Jaro-Winkler algorithm, in order to identify properties owned by the relevant firms (Abood, 2017).

I am able to further identify institutional investor-owned properties by matching owners by common mailing address. Applying this method naively likely produces false positives because the mailing address field in Corelogic is often populated by the address to law offices, banks, government institutions, or other entities that could be involved in a property transaction other than the direct owner. To combat this, I compile a list of search strings for banks, government institutions, law offices, and construction companies, and use the list to identify non-rental owners. I remove these observations from the sample, and remove any associated addresses from the imputation procedure.

4.3 Sample Construction

To ensure that the sample consists of primarily single-family home rentees, I further restrict the sample to absentee-owned properties coded by CoreLogic as single family residences, townhouses, or condominiums. With this restriction I assume that these property types are primary competitors in the SFR market.

I also drop implausible rent price observations in the raw data – specifically where the SFR rent price is under $500 or over $5,000. Rents are assumed to be monthly, which is consistent with the overall distribution of the data.

Zip code demographics data is taken from the ACS5 surveys conducted by the Census Bureau. Zip code level statistics are only available as 5-year intervals, starting from 2007-2011 up until 2014-2018. I map the data from these intervals to specific years, such that the data from each ACS5 release year is allocated as evenly as possible, and is associated with the median of their year coverage.

In this study I assume each zip code to define a regional market. This is primarily out of convenience as the Zillow data is available only at a zip code level of granularity. However, some evidence in the urban economics literature supports the use of zip codes in defining submarkets in terms of their accuracy in delineating market values (Goodman and Thibodeau, 2003). I remove zip codes for which CoreLogic has records for under 50 properties. Finally, based on geographic coverage information provided by Ganduri et al and the websites of the merging firms, the sample is narrowed to the following 10 states: Arizona, California, Florida, Illinois, Georgia, Michigan, North Carolina, Texas, Washington, and Connecticut (Ganduri et al., 2019).

4.4 Measure of Market Concentration

The key variable in my analysis that serves as a measure of local market concentration is the Herfindahl Hirschman Index (HHI) (sum of squared property shares). Use of this index is supported by the Department of Justice in first-glance evaluations of the potential market power gained from proposed M&A activities[9]. Thus, a reasonable measure of the change in market concentration induced by a merger is the change in HHI prior to and following a merger. Unfortunately post-merger market concentration outcomes and prices could be confounded by many potential factors, such as changes in local economic conditions and individual preferences over time. Thus a common measure used to estimate the price effects of mergers is the "simulated change in HHI" ($sim\Delta HHI$), or the anticipated change in HHI that would have occurred following a merger ceterus paribus. The formula for computing this measure is simple

$$sim\Delta HHI = 2 \cdot Firm1Share \cdot Firm2Share \tag{14}$$

This measure allows me to exploit the sharp and heterogeneous increases in local market concentration induced by mergers among the national SFR owners in my sample, while avoiding the endogeneity issue in using market concentration measures directly. Similarly to Ashenfelter et al., I estimate the reduced-form relationship between $sim\Delta HHI$ and rent price in order to get at the causal effect of the increase in market concentration induced by a merger on post-merger prices (Ashenfelter et al., 2015). As noted by Miller and Weinberg, however, the extent of the inferred unilateral price effect captured by these reduced-form estimates is dependent on the extent to which consumer substitution is proportional to market share (Miller and Weinberg, 2017). This assumption may be more reasonable in the housing market than in most other markets, as there are several factors that constrain supply and thus the range of choices consumers have in any local housing market, such as residential space and zoning regulation. Furthermore, I am primarily interested in the sign and relative magnitudes of the price effects, and not the absolute effect sizes.

4.5 Summary Statistics

Table 1 provides some basic summary statistics on the initial set of mergers I choose to study. The property counts for each firm is defined as the number of unique properties for which I observe rent data in my constructed sample. The true pre-merger property counts based off news reports is reported in parentheses. It is clear that my method of identifying merging

[9]https://www.justice.gov/atr/horizontal-merger-guidelines-08192010

firms is not able to consistently identify properties across different firms, and that there is almost certainly some overcount among the properties identified for Beazer Pre-Owned Rental Homes and Invitation Homes, and undercount for the other firms.

There are a number of possible reasons for this discrepancy in property identification. First is that my primary source for firm subsidiaries are filings from the SEC supplemented by a list manually compiled by (Abood, 2017). The public companies in this sample are American Homes 4 Rent, Invitation Homes, and Starwood Waypoint Residential. The rest of the companies are identified solely off of Abood's list and owner address imputation, and property identification is therefore not as reliable. Another factor is that the owner data in CoreLogic do not necessarily correspond to the final owner of the property. As noted above, the owner address and name fields are often populated by details for the law office or mortgage firm involved in the transaction. Due to time and resource constraints, I am unable to further develop the procedure for property identification, and for the sake of my empirical estimates assume that the properties I identify for each firm are drawn from the same distribution as the other properties in my sample. Due to insufficient capture of properties for American Residential Properties and Tricon Capital Group I drop the corresponding mergers from my analysis. Thus the three mergers that I make the final focus of this study are the mergers between Beazer Pre-Owned Rental Homes and American Homes 4 Rent, Colony American Homes and Starwood Waypoint Residential, and Starwood Waypoint Residential and Invitation Homes.

Table 1: Mergers Summary

Target				Acquiror				Announcement Date	Completion Date
Name	Properties	Share	S.D.	Name	Properties	Share	S.D.		
Beazer Pre-Owned Rental Homes	1,404 (1300)	0.58%	0.53	American Homes 4 Rent	8,379 (25,505)	0.34%	0.93	7/1/14	7/1/14
Colony American Homes Inc	3,507 (19,000)	0.12%	0.14	Starwood Waypoint Residential	10,770 (12,500)	0.35%	0.43	9/21/15	1/5/16
American Residential Properties Inc	280 (8,938)	0.07%	0.14	American Homes 4 Rent	15,005 (38,377)	0.41%	0.72	12/3/15	2/29/16
Starwood Waypoint Residential	20,496 (32,000)	0.44%	0.68	Invitation Homes Inc	49,619 (50,000)	0.91%	1.57	8/10/17	11/16/17
Silver Bay Realty Trust Corp	9,640 (9,044)	0.44%	0.18	Tricon Capital Group Inc	0 (7,000)	NA	NA	2/27/17	5/9/17

Notes: The above table summarizes 6 recent mergers of large institutional SFR owners. Properties are counted as the total pre-merger properties with rent from the constructed sample. The pre-merger property counts in parentheses are extracted from news coverage of the respective mergers (`housingwire.com`). HHI and shares data is computed through the constructed sample and represent means of pre-merger values unless otherwise stated. The merger announcement and completion dates are taken from the ThomsonONE database.

Figure 1 shows the identified merging firm's properties plotted across the US. There is a significant degree of sorting of firms across states, as well as geographic clustering within the urban centers of each state, consistent with the firms' stated goals of achieving efficiency through local density. The plots for Arizona and Texas in Figure 2 for instance, shows significant geographic overlap between Invitation Homes and Starwood Waypoint Residential, as well as between American Homes 4 Rent and Beazer Pre-Owned Rental Homes. Similar local overlap of merging properties can be seen in the plots for the other states, which can be found in Appendix E.

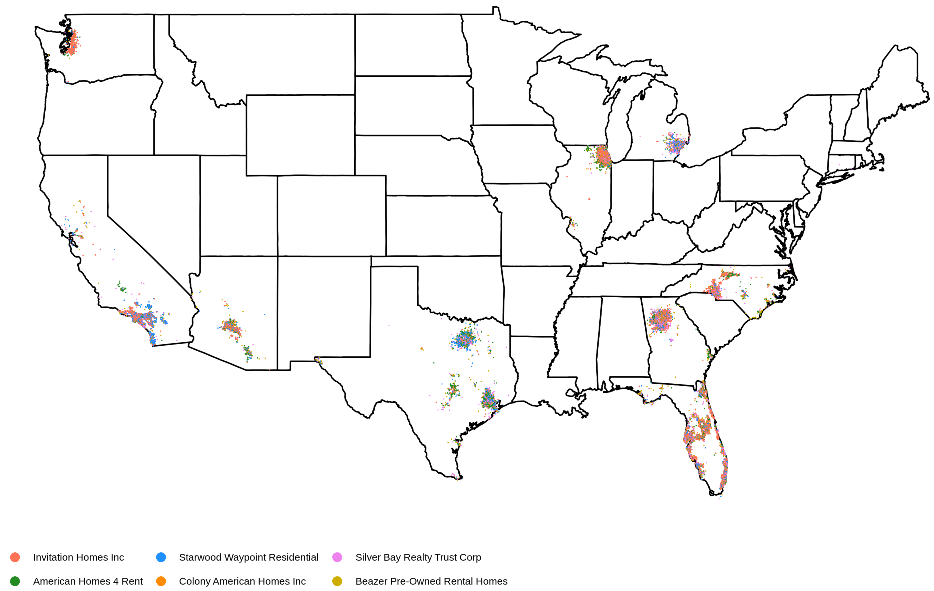

Figure 1: SFR Owners in the US

Figure 2: SFR Owners in TX and AZ

Rent summary statistics are reported in Table 2. Mean rents are recorded for each firm for the pre and post period of each merger, and mean ZORI for the zip codes affected by each merger. For this study I define an affected zip code market as one in which both merging firms are present pre-merger. Standard deviations are reported in parentheses. There is a general upwards trend for rents, as demonstrated by both the ZORI means and the pre and post merger rent means for each merging firm. Unsurprisingly, the distribution of the rent prices for each firm are similar to the distribution of the rents of their regional markets.

Table 2: Rent Summary

Target			Acquiror			Pre-Merger ZORI	Post-Merger ZORI
Name	Pre-Merger	Post-Merger	Name	Pre-Merger	Post-Merger		
Beazer Pre-Owned Rental Homes	1,238	1,366	American Homes 4 Rent	1,362	1,468	1,330	1,533
	(473.48)	(526.68)		(425.99)	(409.79)	(423.22)	(493.54)
Colony American Homes Inc	1,290	1,492	Starwood Waypoint Residential	1,371	1,487	1,214	1,401
	(375.63)	(324.44)		(469.94)	(411.33)	(258.03)	(268.05)
Starwood Waypoint Residential	1,387	1,501	Invitation Homes Inc	1,412	1,561	1,571	1,782
	(465.96)	(391.09)		(506.18)	(483.12)	(639.46)	(703.38)

Notes: The above table summarizes the rent data for each merger in the constructed property sample. The target and acquiror rent values are means of the property-level rent data from MLS, and the ZORI Zillow means are computed across all zip codes affected by each merger. Standard deviations are in parentheses.

Table 3 displays means of pre-and-post merger HHI, quantiles of the predicted increase in concentration variable $sim\Delta HH$, and the total number of zip codes affected by each merger. HHI across all affected zip codes is quite low, and the percentiles show that the distribution of simulated increase in concentration is heavily left-skewed. For context, in the 2010 Horizontal Mergers Guidelines, the DOJ classifies markets into three types based on HHI. All markets with HHI under 1500 are considered unconcentrated, and "any mergers that result in unconcentrated markets are unlikely to have adverse competitive effects"[10].

There is not a single market in my sample which has a post-merger HHI that surpasses the threshold of an unconcentrated market by this definition, though the maximum value of $sim\Delta HH$ for the Starwood–Invitation merger could certainly raise some eyebrows. Of course HHI is a simple heuristic for determining which mergers are the most cause for concern, and does not capture other important factors that determine post-merger welfare, such as the strategic reaction of other firms in the market and possible synergy gains from merging. Furthermore, it is likely that my measures of HHI understate the true level of concentration in each market, as properties recorded under different company names or addresses could still be under the same general management, or be subsidiaries/shell companies of larger property owners. Zip codes could also be too large of an area to adequately capture the level at which property owners compete locally. Nevertheless, it is clear from the above estimates of simulated concentration increase and pre-merger shares that the price change induced by unilateral competitive effects of these mergers should be modest at best.

5 Empirical Strategy and Results

There are two key questions I am interested in answering in this study. (1) What is the price effect of the recently completed mergers between institutional SFR investors, as a result of the induced increase in local market concentration? And (2) is there heterogeneity in these price effects resulting from differences in pre-merger market structures and marginal cost

[10] https://www.justice.gov/atr/horizontal-merger-guidelines-08192010

Table 3: Merging Markets Summary

Merger	Pre-Merger	HHI	$sim\Delta HHI$ Avg	10Pctl	50Pctl	100Pctl	# Zip Code Markets
Beazer Pre-Owned–American Homes 4 Rent	24.95	27.17	0.71	0.01	0.14	34.32	4,172
	(61.49)	(56.95)	(2.02)				
Colony American–Starwood Waypoint	32.20	39.77	0.09	0.005	0.04	0.99	2,512
	(64.43)	(76.34)	(0.12)				
Starwood Waypoint–Invitation Homes	25.27	27.74	2.37	0.03	0.46	179.66	4,504
	(64.83)	(63.30)	(9.33)				

Notes: The above table summarizes the zip code markets I identify in my data that are affected by each of the 3 mergers. HHI data is computed through the constructed sample and represent means across all merge-affected zip codes. 10Pctl, 50Pctl, and 100Pctl represent the 10th, 50th, and 100th percentile values of $sim\Delta HHI$ for each merger. Standard deviation is in parentheses.

outcomes, as suggested by the model in Section 3? Public communications surrounding the SFR mergers primarily emphasize the operational efficiency that would be captured through greater geographic density of property ownership[11]. I am interested in evaluating to what extent these claims of efficiency are supported by the data. The granularity of the dataset will allow me to estimate separate price effects for each of the affected market participants in a merger. Specifically, I use a 2-way fixed effects model that exploits the variation in prices within and across different local markets over time, as well as variation in pre-merger market shares, in order to identify the aforementioned effects.

The key identifying assumption of my approach is that there are no time-varying market-specific factors that are correlated with both pre-merger market shares and rent prices, along with the standard parallel trends assumption for difference-in-differences designs. I discuss the extent to which these assumptions hold in the data and how violations of these assumptions could affect the effects I estimate in Section 6.1.

As discussed in Section 4.4, I rely on the predicted increase in concentration $sim\Delta HHI$ to estimate the causal effect of merger-induced concentration on prices. In order to confirm that $sim\Delta HHI$ successfully predicts post-merger HHI, I estimate the following equation:

$$HHI_{zt} = \beta sim\Delta HHI_z * \text{Post}_{zt} + \alpha_t + \gamma_z + [X_{zt}] + [\gamma_z * t] + [sim\Delta HHI_z * t] + \epsilon_{zt} \quad (15)$$

where the dependent variable is the computed HHI for zip code z in year t, X_{zt} contains zip-code level controls, and α_t and γ_z are year and zip fixed effects, respectively. Note that I estimate this equation on the sample from 2009-2018 containing all mergers (5 years prior to the first merger), so the post-period varies by zip code depending on which merger is in effect, thus the z subscript on the $Post$ dummy. To ensure robustness of the relationship, I estimate the equation controlling for time-varying zip-level unemployment rates and median household income (X_{zt}), zip-code linear time trends ($\gamma_z * t$), and predicted change in concentration

[11]https://www.sec.gov/Archives/edgar/data/0001687229/000119312517348234/d494074d8k.htm
https://www.sec.gov/Archives/edgar/data/1562401/000119312514261977/d753436dex21.htm
https://www.sec.gov/Archives/edgar/data/1579471/000119312515406507/d107500dex991.htm

interacted with a linear time trend ($sim\Delta HHI_z * t$). Table 3 shows the results. As predicted, $sim\Delta HHI$ is a strong predictor of post-merger HHI, and this relationship is robust across the full set of controls.

Table 4: Post-Merger Concentration on Simulated Concentration

	Dependent variable:			
	HHI			
	(1)	(2)	(3)	(4)
$sim\Delta HHI \times$ Post	1.693***	1.692***	0.983***	1.907***
	(0.167)	(0.129)	(0.089)	(0.274)
Zip Controls	No	Yes	Yes	Yes
Zip Time Trend	No	No	Yes	No
$sim\Delta HHI$ Time Trend	No	No	No	Yes
Observations	15,334	15,314	15,314	15,314
R^2	0.957	0.957	0.985	0.957
Note:		*p<0.1; **p<0.05; ***p<0.01		

5.1 Effect of Mergers on Zip-Code Rents

I estimate the effect of the mergers on overall market prices using the Zillow ZORI data. Specifically I estimate the following regression:

$$\begin{aligned} \log P_{zt} = &\beta sim\Delta HHI_z * \text{Post}_{zt} + \alpha_t + \gamma_z + [X_{zt}] + [\gamma_z * t] + \\ &[sim\Delta HHI_z * t] + \epsilon_{zt} \end{aligned} \tag{16}$$

where all the controls are the same as in (15) except that the covariate vector X_{zt}, now includes median property tax value and median property log square-footage on top of the original zip code controls. I estimate (16) on the pooled mergers data as well as for each individual merger, where the control group consists of zip codes for which exactly one of the merging firms owns properties in the pre-merger period.

Table 8 shows the resulting estimated coefficient on $sim\Delta HHI_z * \text{Post}_{zt}$ for each sample, where "Pooled" denotes the total pooled mergers sample. "OLS" contains the estimates on the basic 2-way FE regression with no controls, "Property Controls" includes controls for median property characteristics, "ACS Controls" includes controls for zip code characteristics, "Zip Trend" includes zip time trends along with the previous controls, and "$sim\Delta HHI$ Trend" includes $sim\Delta HHI$ time trends along with the property and ACS controls. Standard errors are in parentheses.

The estimates for the Beazer–American Homes 4 Rent (AH4) merger and Starwood–Invitation Homes merger are significant across the full range of controls, with their effect size somewhat

Table 5: Effect of Simulated Concentration on Zip Code Prices

	Coefficient: $sim\Delta HHI * Post$				
	OLS	Property Controls	ACS Controls	Zip Trend	$sim\Delta HHI$ Trend
Pooled	0.001^{***}	0.001^{***}	0.0004^{***}	0.0004^{***}	-0.002^{***}
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.001)
	[131,580]	[131,580]	[108,885]	[108,884]	[108,885]
Beazer – AH4	-0.003^{***}	-0.002^{***}	-0.002^{***}	-0.002^{***}	-0.002^{***}
	(0.001)	(0.001)	(0.001)	(0.001)	(0.0005)
	[127,338]	[127,338]	[105,408]	[105,408]	[105,408]
Colony – Starwood	-0.006	0.0005	-0.011	-0.011	-0.018^{***}
	(0.013)	(0.013)	(0.010)	(0.010)	(0.006)
	[128,017]	[128,017]	[105,960]	[105,960]	[105,960]
Starwood – Invitation	0.001^{***}	0.001^{***}	0.0005^{***}	0.0005^{***}	0.0004^{***}
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
	[130,748]	[130,748]	[108,205]	[108,204]	[108,205]

Notes: The table reports estimates of the coefficient β in (16) for each merger and the total pooled sample. *, **, *** indicate significance at the 0.1, 0.05, and 0.01 levels respectively. Standard errors for the coefficients are reported in parentheses, and sample sizes are reported in brackets. The average $sim\Delta HHI$ are 0.71, 0.09, and 2.37, for the Beazer–AH4, Colony–Starwood, and Starwood–Invitation mergers, respectively.

attenuated with the inclusion of additional controls. The Colony-Starwood estimates are insignificant until controlling for the $sim\Delta HHI$ trend, which may be a sign that the trend control is a poor fit. The sign of the estimate for the pooled sample flips with the inclusion of the $sim\Delta HHI$ trend, which may be of some concern. Given that the relationship between predicted concentration and post-merger rent have different signs between the Beazer-AH4 and Starwood-Invitation mergers, this could be a consequence of one trend dominating the other in the pooled sample.

To investigate this further, I estimate a more flexible version of (16) that includes interactions of simulated concentration increase with each individual month relative to the merger. Specifically, I estimate the following with the same set of property and zip code controls as in (16), restricting the sample period to 30 months prior to or 40 months after the merger.

$$\log P_{zt} = \sum_{j=-30}^{40} \beta_j sim\Delta HHI_z * 1(\tau_t = j) + X_{zt} + \alpha_t + \gamma_z + \epsilon_{zt} \tag{17}$$

Here $1(\tau_t = j)$ is a year-month dummy for the month relative to merger completion. I normalize β_0 to 0, and plot the estimated coefficients β_j in Figure 3, scaling each coefficient by average simulated increase in HHI for each merger, with associated standard error ribbons.

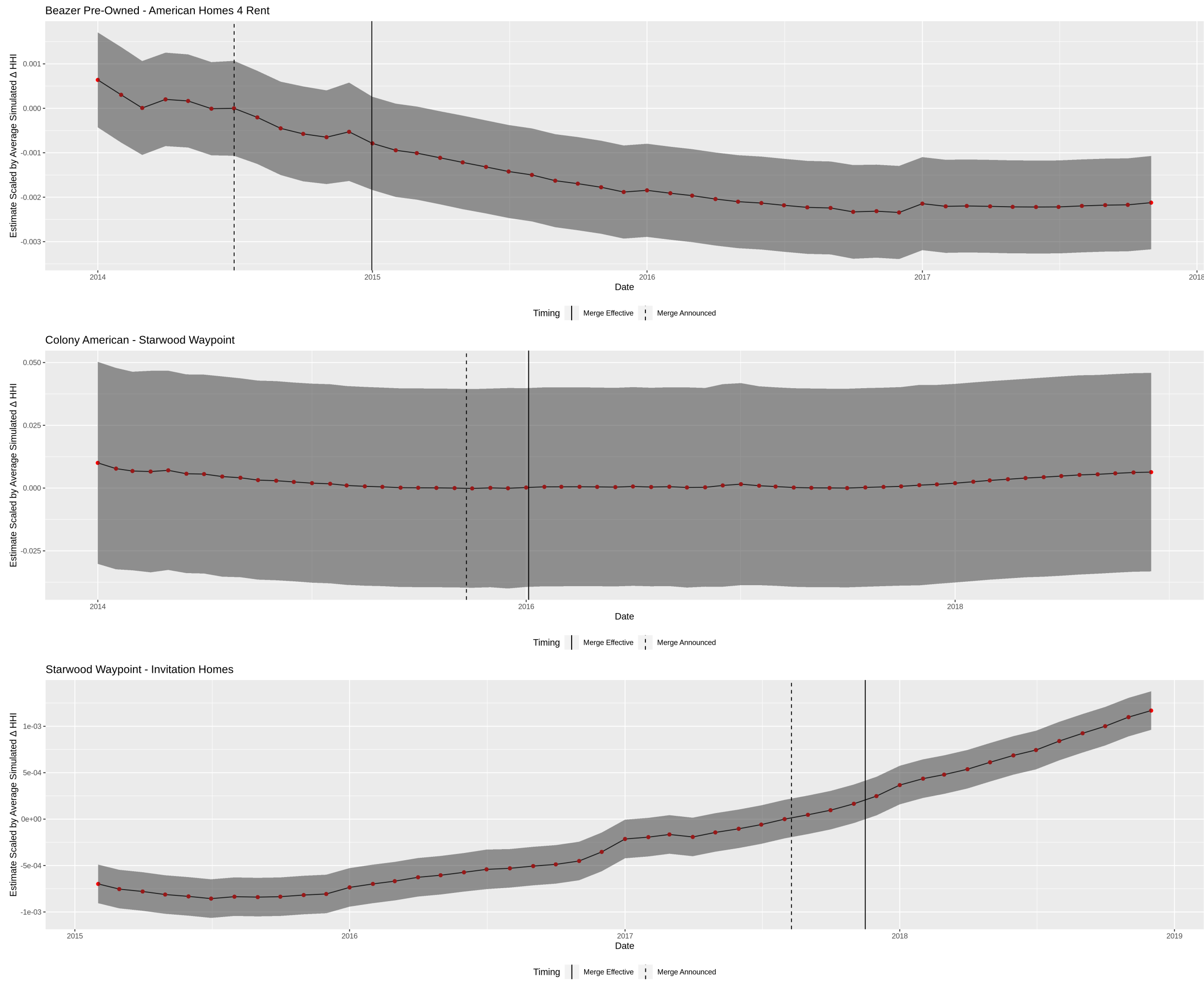

Figure 3: Coefficients on Year/Month Interacted with $sim\Delta HHI$

There are a number of important implications we can draw from Figure 3. First, it is clear that there are significant pre-trends associated with the Starwood-Invitation merger, where zip codes with higher pre-merger shares start off with relatively low rent, and experience a disproportionate growth in rent price starting from the end of 2015. On the other hand, the Beazer-AH4 merger coefficients show the reverse trend, where higher share zips start with relatively high rent, and experience a steady fall in rents over time. The pre-period for the Beazer-AH4 case is around 7 months, and thus not long enough to say for certain that the suggested trend is present under a longer pre-period[12].

The slope of the curves helps to resolve the issue of the sign flip when controlling for the $sim\Delta HHI$ time trend in the pooled estimates, when considered alongisde a number of features of my data. Table 1 and Table 3 show that the Starwood–Invitation merger makes up the majority of the treated observations in my sample, and that $sim\Delta HHI$ is much higher in general for zip codes affected by this merger. Furthermore, the trend curve for the

[12]I show in my property-level estimates that the trend is persistent for up to 5 years pre-merger

Starwood–Invitation merger in Figure 8 is steeper than that for the Beazer–AH4 merger. These facts result in the coefficient on the $sim\Delta HHI$ time trend control to be positive[13], which controls for the positive price effect among the Starwood–Invitation zips, but consequently further biases downwards the negative estimates from the Beazer-AH4 merger. Given these stark differences in direction of effect and price trend across the three mergers, it is clear that attempting to derive some estimate of a "general" price effect of SFR merger-induced concentration change would disregard important merger-specific factors that drive this observed effect heterogeneity and pool these offsetting effects, resulting in a weaker and less-informative estimate.

The pre-trends uncovered in Figure 3 illustrate the importance of controlling for the $sim\Delta HHI$ time trend, and boost confidence that the estimates in the last column of Table 8 reflect the true price effects induced by the Beazer–AH4 and Starwood–Invitation mergers. The large standard errors and slight convexity of the trend curve for the Colony–Starwood merger suggest that the negative effect found only in the last column is a result of a poor fit of the linear $sim\Delta HHI$ trend.

Taking the average simulated increase in concentration for each merger from Table 3, the estimates in Table 8 suggest that the Beazer–AH4 merger caused an average rent decrease of around .14%, while the Starwood–Invitation merger caused an average rent increase of around .11%. Using the maximum values of $sim\Delta HHI$, the upper bound of price change for a local market in this sample would be a rent decrease and increase both around 7%. As predicted in the previous section, the effects of these mergers on average equilibrium rent prices are very small to be almost economically negligible, and even in the most concentrated markets in the sample result in an average price change of under 10%. I also provide estimates of (16) excluding the interaction with $sim\Delta HHI$ in Appendix B, and the estimated effect sizes are similar to the coefficients in Table 7 scaled by $sim\Delta HHI$, confirming the relationship between merger-induced concentration and market prices.

5.2 Heterogeneity in Price Effects

Though the zip-level estimates suggest that the mergers did not have much of a effect on average post-merger market prices, there are a number of outstanding questions and concerns that remain to be addressed. One possible concern is that the Zillow data starts in 2014, and therefore there may not be a long enough post-period to accurately capture the counterfactual trend for the Beazer–AH4 merger. Another is that the Zillow data is only available for the top 100 major metropolitan areas in the United States, and thus a significant portion of the treated population is left out of the estimation. For context, the property-level sample covers 18%, 17%, and 20% more zip codes than Zillow for the Beazer–AH4, Colony–Starwood, and Starwood–Invitation mergers, respectively. Finally, the zip-level estimates are unable to address the possible heterogeneity in price effect postulated by the 3-firm model in Section 3, depending on the degree of merger-induced efficiencies along with other pre-merger market factors.

In order to address these concerns, I estimate the below equation using the sample con-

[13] I confirm during estimation that the estimated coefficient is positive – specifically 0.00005

structed using the Corelogic property and MLS prices data:

$$
\begin{aligned}
\log P_{izt} =& \beta_O sim\Delta HHI_z * \text{Post}_{zt} * \text{Outside}_{it} + \\
& \beta_A sim\Delta HHI_z * \text{Post}_{zt} * \text{Acquiror}_{it} + \\
& \beta_T sim\Delta HHI_z * \text{Post}_{zt} * \text{Target}_{it} + \\
& \alpha_t + \gamma_z + + X_{izt} + [\gamma_z * t] + [sim\Delta HHI_z * t] + \epsilon_{izt}
\end{aligned} \tag{18}
$$

where “Outside”, “Acquiror”, and “Target” are dummy indicators for whether a treated property belongs to an outside firm, the acquiring firm, or the target firm, respectively. X_{izt} contains controls for log square footage and tax value for each property, and unemployment rate and median household income for each zip. I estimate (18) on the three mergers and present the estimates of β_O, β_A, and β_T in Table 9, for varying levels of controls. As with the previous section, my control group consists of all properties in the zip codes which have exactly one of the merging firms present pre-merger.

Table 6: Effect of Simulated Concentration on Property Prices

	Property/Zip Controls			Zip Trend			$sim\Delta HHI$ Trend		
	Outside	Target	Acquiror	Outside	Target	Acquiror	Outside	Target	Acquiror
Beazer - AH4	−0.003***	−0.004***	−0.001	0.003***	0.002	0.004***	0.003***	0.002*	0.004***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
		[2,825,812]			[2,825,810]			[2,825,812]	
Colony - Starwood	0.051**	−0.189**	−0.029	0.149***	−0.101	0.094***	0.150***	−0.091	0.072**
	(0.022)	(0.073)	(0.030)	(0.023)	(0.081)	(0.031)	(0.024)	(0.074)	(0.030)
		[2,604,538]			[2,604,536]			[2,604,538]	
Starwood - Invitation	0.0008***	0.0009***	0.0008***	0.0008***	0.001***	0.0008***	0.0008***	0.0009***	0.0008***
	(0.0002)	(0.0003)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
		[2,150,581]			[2,150,577]			[2,150,581]	

Notes: The table reports estimates of the coefficients β_O, β_A, and β_T in (16) for each merger. *, **, *** indicate significance at the 0.1, 0.05, and 0.01 levels respectively. Standard errors for the coefficients are reported in parentheses, and sample sizes are reported in brackets. The average $sim\Delta HHI$ are 0.75, 0.01, and 3.38, for the Beazer–AH4, Colony–Starwood, and Starwood–Invitation mergers, respectively.

Here “Outside”, “Target”, and “Acquiror” refer to the estimates of β_O, β_A, and β_T, respectively. “Zip Trend” and “$sim\Delta HHI$ Trend” report the estimates controlling for property/zip characteristics and either zip code time trends or $sim\Delta HHI$ time trends, respectively. Standard errors are reported in parentheses, and sample sizes are reported in brackets.

There is significant instability in the property-level regressions when controlling for zip code and $sim\Delta HHI$ time trends. The fact that all the coefficients for the Beazer–AH4 and Colony–Starwood mergers are pushed up by similar magnitudes when controlling for time trends suggests that linear trends do not fit the shape of the data well. To investigate this for the zip-code time trends, I plot the residuals from a 2-way FE regression of log rent with zip linear time trends in Figure 4. The unweighted means of the residuals are very close to zero for all the mergers, whereas the means weighted by $sim\Delta HHI$ are convex, signaling the need to control for quadratic time trends when estimating 18 with the $sim\Delta HHI$ interaction.

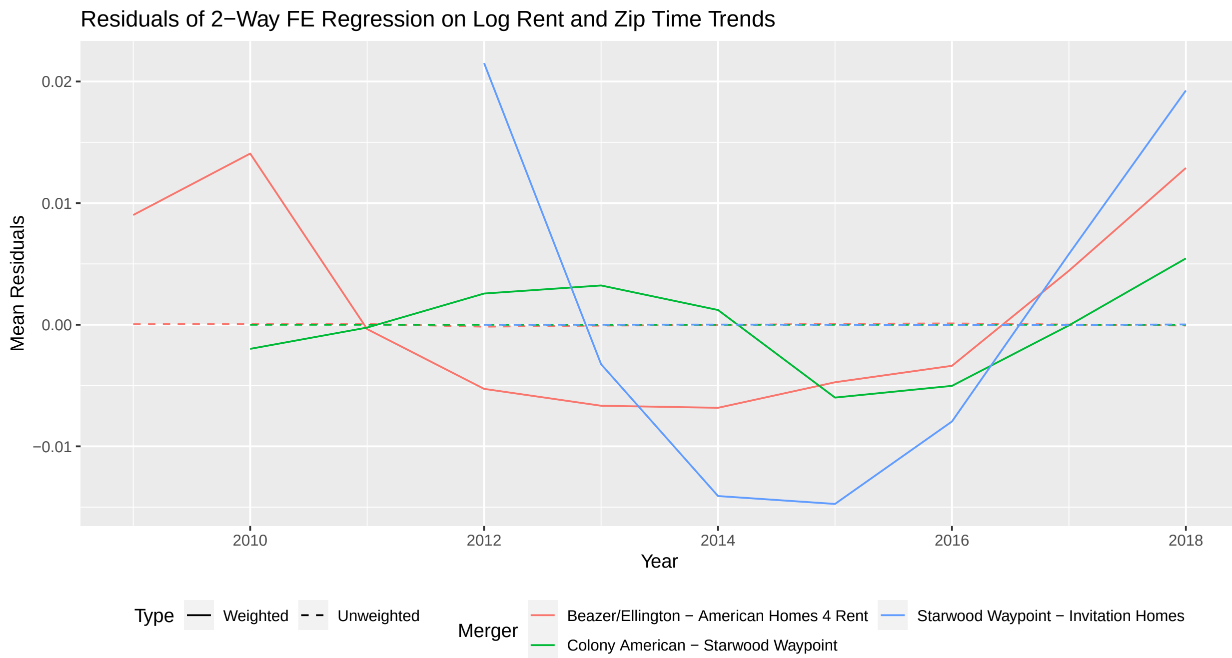

Figure 4: Residuals for 2-Way FE Rent Regression Controlling for Zip Linear Time Trends

I follow up on the $sim\Delta HHI$ time trends by estimating a version of (17) except on years and individual properties, with the sample restricted to 5 years before and after each merger.

$$\log P_{izt} = \sum_{j=-5}^{5} \beta_j sim\Delta HHI_z * 1(\tau_t = j) + X_{izt} + \alpha_t + \gamma_z + \epsilon_{izt} \tag{19}$$

Figure 5 plots the event-time coefficients scaled by average $sim\Delta HHI$ with standard error ribbons. It is clear that with the extension of the pre-period, there is now significant convexity in the relationship between pre-merger shares and rent across all mergers.

Figure 5 also highlights a significant discrepancy in the post-period trend for the Beazer–AH4 merger between the Zillow and MLS datasets. The post-period curve in Figure 3 showed a steady decline into 2018, whereas the property-level plot remains above 0 and begins trending upwards after 2016. The trend shape remains consistent even when I estimate the event-time regression on just the zip codes present in the Zillow data, which indicates that differences in the underlying rent data is driving this discrepancy. To verify this, I plot the difference in median zip rent trend between the MLS and Zillow data, weighted and unweighted by $sim\Delta HHI$. Figure 6 shows that zip codes with higher pre-merger firm shares also happen to have higher reported rent prices in the MLS data relative to the Zillow data. This could result from some systematic bias in the MLS databases that Corelogic sources from, an issue in my sample construction procedure, or something else entirely. Due to time constraints, I am unable to conduct a deeper investigation into this issue. Given that Zillow's data is built

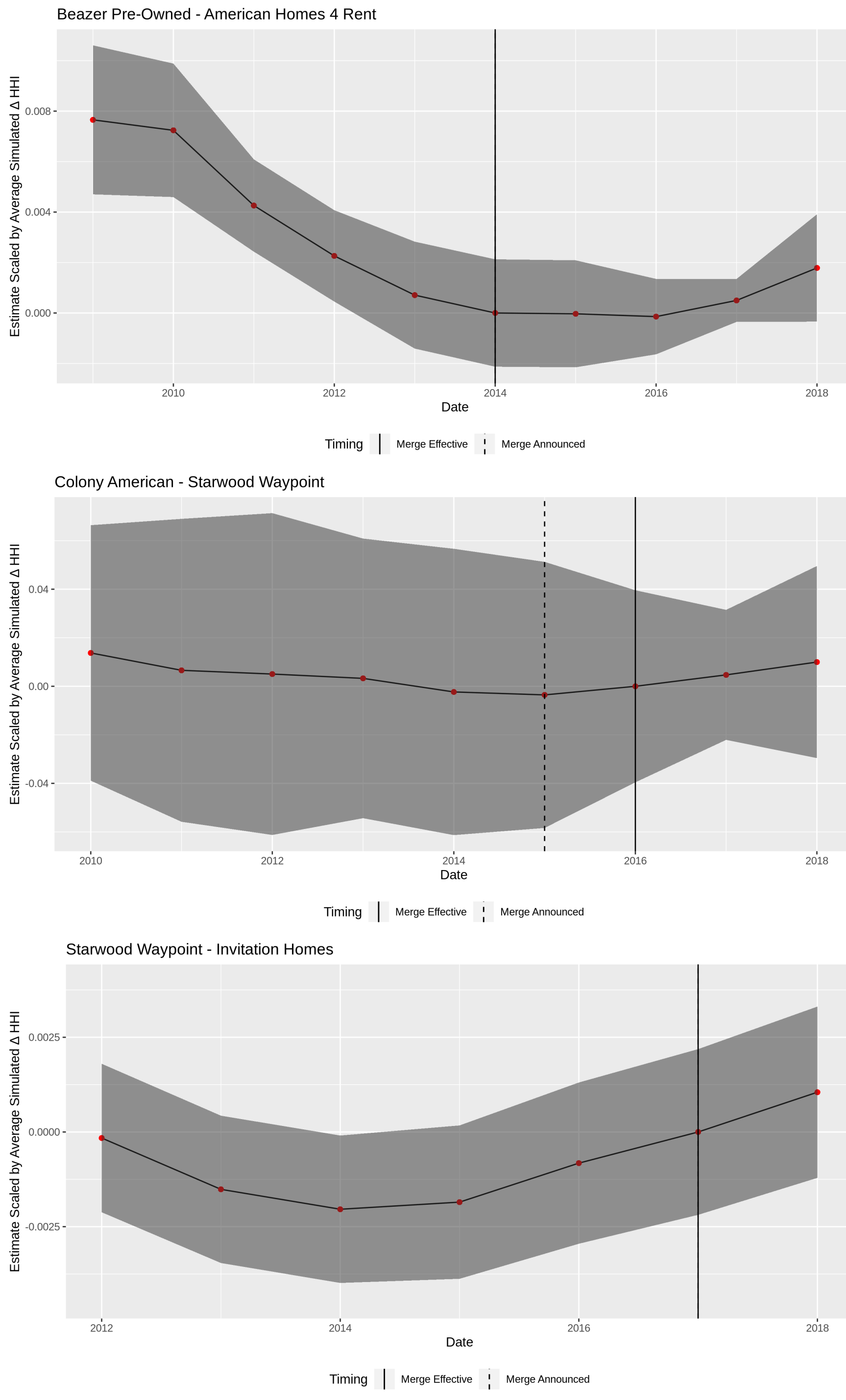

Figure 5: Coefficients on Year Interacted with $sim\Delta HHI$

from its own record of home listings and thoroughly cleaned/vetted by its internal team, however, the zip code estimates are likely more reliable in capturing the true general price effects of the Beazer–AH4 merger.

Given the convexity of the curves in Figure 5 and shape of the residuals in Figure 4, I re-

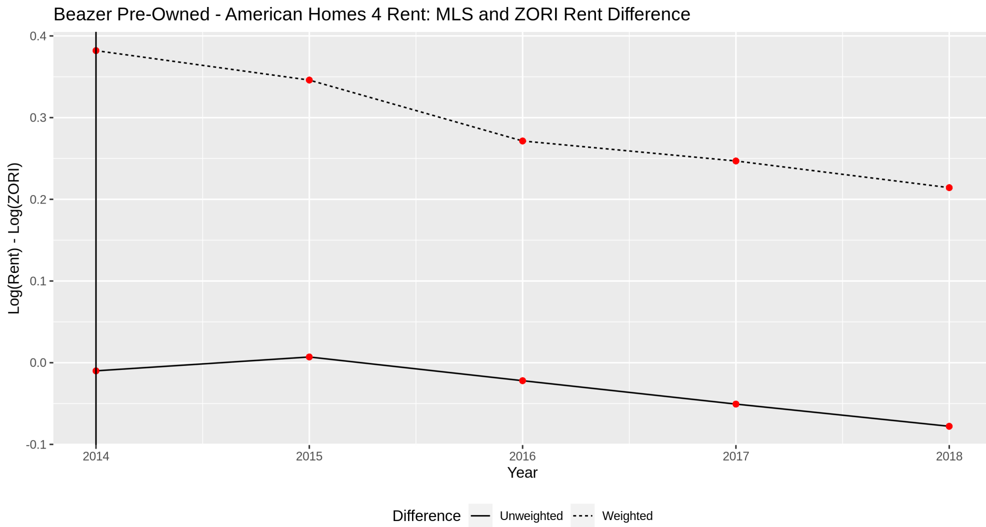

Figure 6: Difference in MLS and Zillow Rent Trends for Beazer–AH4 Affected Zips

estimate (18) controlling for both linear and quadratic time trends. The coefficients for these estimates are in Table 7.

I highlight in bold the estimates that are robust across the set of controls. As expected the positive bias in the $sim\Delta HHI$ coefficients goes away once the trend is properly controlled for. For the Starwood–Invitation merger, the estimates are slightly more attenuated compared to the zip code estimates. In the Beazer–AH4 case, controlling for quadratic trends causes the estimates to lose their significance entirely. This result, coupled with the fact that the MLS rent data for the affected zip codes for this merger is of suspect quality, means that we are unable to say anything about heterogeneity in the price effects of the Beazer–AH4 merger.

With the other two mergers, however, I am able to identify some interesting heterogeneity in the price outcomes of the merging market participants. The Colony–Starwood estimates are especially intriguing because they suggest that the lack of effect found in the zip code estimates are a result of opposing price changes between outside and target properties. Specifically, outsider properties within the markets affected by the Colony–Starwood merger experienced a slight increase in rent while the merging Starwood Waypoint properties experienced a slight decrease. These offsetting price changes are consistent with the null effect found for the zip-code level estimates. With the Starwood–Invitation merger, all properties within a local market appear to increase in rent, with the acquiring Invitation properties experiencing around half the percentage increase of the other properties.

Based on my estimates and pre-merger market shares, Colony American Homes properties (the target in the Colony–Starwood merger) experienced an average rent decrease of 0.2% with an upper bound (based on the maximum $sim\Delta HHI$ in my sample) decrease of 1.9%, while outsider competing properties experienced an average rent increase of 0.5% with an upper bound increase of 5%. For the Starwood–Invitation case, both the target (Starwood Waypoint) and outsider competing properties had an average .17% increase with an upper bound of 9%, while the acquiror (Invitation Homes) properties had an average .07% increase with an upper bound of 3.6%. These price effects are quite marginal as a result of the low pre-merger shares of SFR firms in my sample, though may still reflect real responses to the change in market structure induced by these mergers in the few zips that have non-trivial pre-merger firm shares.

I also provide estimates of (18) excluding the interaction with $sim\Delta HHI$ in Appendix B, and the estimated effect sizes are similar to the coefficients in Table 7 scaled by $sim\Delta HHI$, confirming the relationship between merger-induced concentration and heterogeneous price change. In the next section I make some predictions regarding the marginal cost changes induced by these mergers by comparing the relative magnitude and direction of these estimates.

Table 7: Effect of Simulated Concentration on Property Prices with Quadratic Time Trends

	Polynomial Zip Trend			Polynomial $sim\Delta HHI$ Trend		
	Outside	Target	Acquiror	Outside	Target	Acquiror
Beazer - AH4	-0.0002	-0.0021***	0.0012	0.0002	−0.0007	0.0015*
	(0.0003)	(0.0007)	(0.0009)	(0.0004)	(0.0008)	(0.0009)
		[2,825,810]			[2,825,812]	
Colony - Starwood	**0.0488***	**−0.1822***	-0.0024	**0.0508***	**−0.1885**	−0.0324
	(0.0099)	(0.0845)	(0.0276)	(0.0105)	(0.0816)	(0.0334)
		[2,604,536]			[2,604,538]	
Starwood - Invitation	**0.0005***	**0.0005***	**0.0002**	**0.0004***	**0.0005***	**0.0002**
	(8.97×10^{-5})	(0.0001)	(8×10^{-5})	(9.7×10^{-5})	(0.0001)	(9×10^{-5})
		[2,150,577]			[2,150,581]	

Notes: The table reports estimates of the coefficients β_O, β_A, and β_T in (16) for each merger, controlling for both linear and quadratic time trend interactions. *, **, *** indicate significance at the 0.1, 0.05, and 0.01 levels respectively. Standard errors for the coefficients are reported in parentheses, and sample sizes are reported in brackets. The average $sim\Delta HHI$ are 0.75, 0.01, and 3.38, for the Beazer–AH4, Colony–Starwood, and Starwood–Invitation mergers, respectively.

5.3 Implications for Marginal Cost Efficiencies

Given my observed empirical estimates of the price effect of each merger, I am able to apply the inequalities derived in (9)-(11) to provide some intuition on the extent of the marginal cost efficiencies generated by each merger.

For the Colony–Starwood merger, we observe the following direction of price change

$$\frac{\Delta p_T}{\Delta p_T} < 0$$
$$\frac{\Delta \pi_O}{\Delta p_O} > 0$$

where T and O indicate the target and outsider firms, respectively. If we assume that all properties are equally substitutable, i.e. the diversion ratios between the outsider, target, and acquiror properties are all equal to 1/2, the pre-merger profit margins for the two merging firms are the same, and take as given the pre-merger average values reported in Table 2, such that the price ratio $p_A/p_T \approx 1$ and quantity demanded ratio $Q_A/Q_T \approx 3$, then after some algebra we are left with the below inequality

$$-(p_A - c'_A) < \Delta c'_T < 0.63\Delta c'_A + \frac{p_T - c'_T}{5} - (p_A - c'_A)$$

As the degree of substitutability between the merging properties increases (or conversely decreases for outsider properties), the necessary decrease in marginal cost to observe the direction of price changes we see grows. On the other hand, if the merger created greater inefficiencies, such that the change in marginal cost of operating the acquiring firm's original properties actually increases, then there need not be any realized efficiencies for the target properties at all. However, if I include into the system the fact that I do not estimate any significant price change for the acquiror properties (negative coefficient with a 1 SE increase above 0), then I get the following relation

$$\begin{gathered}\frac{\Delta p_A}{\Delta p_A} \leq 0 \\ \Rightarrow \Delta c'_T < -\frac{7}{27}(2p_A - c_A)\end{gathered}$$

so that the target firm must not only realize cost efficiencies, but that those efficiencies must be over 50% of the acquiring firm's original profit margin. This is under an assumption of equal diversion ratios between all property types, whereas in reality the properties between two merging firms may be closer substitutes, which would drive the above inequality even lower. Thus the conservative predictions from my model are that the marginal cost savings were realized for the target firm's properties, and were larger in magnitude than the cost savings, if any, realized for the acquiring firm's properties. This scenario could easily take place in SFRs if the target firm's professional network and management platform were not as well-developed as the acquiring firm's, so that operating cost of the acquired properties reduces to the same level as the broader portfolio.

Next I examine the implications from my estimates on the Starwood – Invitation merger. Here the main result is that the target and outside properties experience higher rent growth than the acquiring firm properties. As with the previous merger, I assume equal diversion ratios, same pre-merger price-cost margins between the merging firms, $p_T/p_A \approx 0$, and $Q_A/Q_T \approx 2$, as per Table 2 and Table 1. I do not make any assumptions on the outside firm other than on the pre-merger quantity ratio, Q_T/Q_O, which I assume to be 1. Then from the observed magnitudes of price change

$$\frac{\Delta p_T}{p_T}, \frac{\Delta p_O}{p_O} > \frac{\Delta p_A}{p_A} > 0$$

I am able to derive the following inequalities:

$$
\begin{aligned}
42\Delta c'_T &> 76\Delta c'_A + 4c_A + p_A + 17p_O + 26c_O \\
17\Delta c'_1 + 5\Delta c'_2 &> 7c_2 + 4c_1 - 11p_1
\end{aligned}
$$

where as the pre-merger ratio between target and outside properties Q_T/Q_O decreases, the right-hand side falls. The first inequality implies that given the relative magnitude of price changes and aforementioned assumptions, either the marginal cost of the acquiror firm declines substantially relative to the target firm, or conversely the marginal cost of the target firm increases substantially relative to the acquiring firm. The second inequality, however, also places a lower bound on the possible extent of the marginal cost decline experienced by the merging firms. Specifically, if pre-merger profit margins are relatively small, then the firms must either experience no change or a net increase in marginal cost.

I initiate a very preliminary empirical evaluation of the above marginal cost prediction for the Starwood – Invitation merger by plotting the per-home operating expenses for each firm in Figure 7, using data extracted from the firms' SEC filings. The merger was completed in the 4th quarter, with the sharp spike in reported costs attributable to the "substantial...costs in connection with the Mergers."[14]. The post-merger trend of operating expense maintains the slow but steady upwards trend from before the merger, so it appears that the purported efficiency gains from the merger have not yet materialized in nearly 3 years post-merger. Furthermore, the pre-merger per-home operating cost for Starwood Waypoint was significantly lower than that of Invitation Homes. If I assume that post-merger marginal costs for each of the properties in the merged portfolio are around the same, then this implies an almost 4x increase in operating expenses for the Starwood Waypoint properties, while the operating expenses for the Invitation Homes properties do not change much. The 4-fold increase in marginal cost is an unrealistic figure, so either there is some accounting difference I miss during my data collection or my prior assumption should be taken with severe skepticism.

[14] https://www.sec.gov/Archives/edgar/data/0001687229/000168722918000018/a12312017ihallfs10k.htm

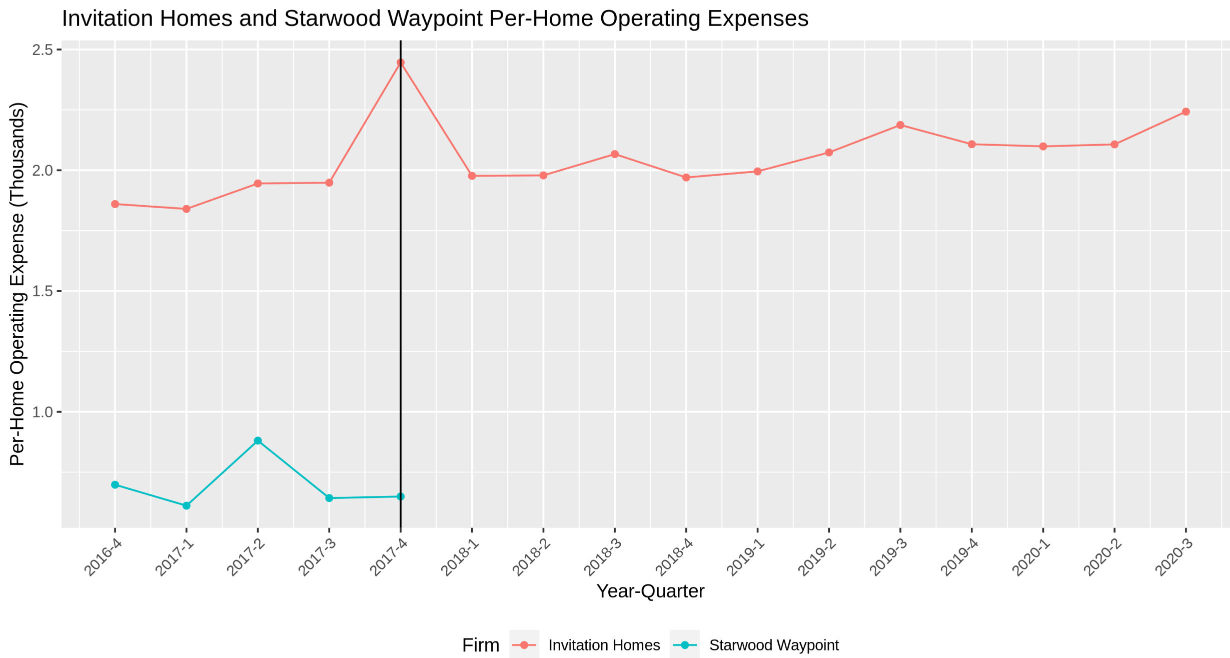

Figure 7: Quarterly Operating Expenses for Invitation Homes and Starwood Waypoint Residential

Nevertheless, setting $\Delta c'_T = 1,500$, $\Delta c'_A = 0$, $c_{2\cdot} = 2,000$, $c_1 = 500$, $p_2, p_1 = 1400$, based on the above results, the inequality computes to

$$63,000 > 46,200$$

Thus this differential increase in marginal cost is consistent with the predictions of the model, and is weak evidence that this increase in property operating expense is what drives the differential increase in rent in the post-period. Note that this does not necessarily mean that the firm experienced a decline in operating efficiency, as there are other factors that could affect both property operating expenses and rent prices other than firm operating efficiency, such as improvement in services, renovations, and similar actions involving product repositioning.

Given the extent of the assumptions I make in both constructing and evaluating my simple model, and the very low effect size of my estimates, the above results should only be seen as suggestive evidence in support of the marginal cost mechanism of merger-induced price change. If the mergers resulted in changes to competitor strategy or neighborhood preferences, then these would not be accounted for in the model. Furthermore, the fact that the housing stock takes time to adjust and markets are constrained by geographic and zoning factors, means that there are substantial frictions to supply adjustment that are not accounted for in my analysis, as I assume equilibrium in the post-merger period. Since I observe a minimum post-period length of 2 years, however, I don't think this is as much of a concern.

6 Discussion

6.1 Parallel Trends Assumption

My identification strategy is subject to a number of key assumptions that are supported by the data to varying degrees. For both my zip-level and property-level estimates, identification of the unilateral price effects of merger induced changes in concentration relies on there being parallel trends in rents between my comparison groups. In the zip code case, this means not only parallel trends between the treated and control (single firm) zip codes, but also parallel rent trends between the merge affected zip codes for different levels of predicted increase in concentration. The property-level case is even more strict with identification of heterogeneous price effects, where I assume parallel trends between the outside, target, and acquiror owned properties within each merge-affected zip.

I present some graphical evidence of parallel pre-trends between the quartiles of simulated concentration increase for the Zillow data in Figure 8 and Figure 9 in Appendix A. These trends look to be mostly parallel, though the Beazer–AH4 merger shows the same decreasing pre-trend as observed in the results, and the Starwood–Invitation pretrend for Q4 zips appears to be steeper than the rest. Given how small my estimated effects are, these pre-trend differences could certainly be driving the effect I find in my zip code estimates for these two mergers. These pre-trend differences are far more substantial in the property-level data as a result of my extended pre-period. On top of the estimates in Figure 5, I estimate another version of (19) with dummies for firm ownership, and plot them in Figure 10 in Appendix A. Though very imprecise, the trends plotted for each of the market participants roughly follow those observed in Figure 5.

I account for these pre-trend differences across zips somewhat by the inclusion of zip code and $sim\Delta HHI$ time trends in my estimates. The stability of my estimates when including these controls for the Colony–Starwood and Starwood–Invitation mergers lends credence to their validity. If these controls are insufficient for dealing with the pre-trends I observe, however, then these trends could still be driving the significance of my estimates.

As emphasized by (Kahn-Lang and Lang, 2020), even parallel pre-trends are not sufficient for proving parallel counterfactual trends. Furthermore, the initial gap in rent price between the treated and control zips may indicate the presence of differences in underlying market characteristics which could drive differences in response to changes in concentration.

In order to address this, I also plot the difference in trends for various property and zip code characteristics in Appendix A. These figures demonstrate the comparability in zip codes across different levels of pre-merger firm shares, for most of the characteristics I record. The differential dip in median household income and property tax values and spike in unemployment for control zips for the Starwood–Invitation merger in Figure 11-Figure 13 may be some cause for concern. For the control group, however, $sim\Delta HHI = 0$ by definition, so these zips only affect the fixed effect estimates for the time dummies in my model.

The degree to which my estimates can be interpreted as capturing the unilateral price effects of mergers is dependent on the extent to which consumer substitution is proportional to market share. This assumption is supported by the equivalence between my estimates with and without interacting with $sim\Delta HHI$, which I report in Appendix B, as well as by the particular institutional setting. Local housing markets will always be supply-constrained as a consequence of their geographic boundaries, and thus consumers who choose to live in a neighborhood with concentrated ownership will necessarily have to choose between one of a few owners.

6.2 Future Extensions

As discussed in the previous section, the empirical estimation I perform in this paper are subject to a number of important assumptions that are modestly supported by the data. On top of these possible threats to identification, my estimates are limited in that they do not reveal the overall welfare effects of past SFR mergers and are not helpful in making predictions about the effects of future mergers among institutional SFR owners. Therefore, there is good motivation to pursue estimation of a structural demand system. Estimation of such a model would require, on top of the micro property data used for this paper, data on individual residential histories, along with individual characteristics or estimated distributions. Data that is robust enough to perform this estimation may be difficult to come by, but is certainly not impossible – similar structural estimation of demand for housing has already been done using the restricted-use micro demographic data from the Census Bureau, for instance, by Bayer et al. (Bayer et al., 2004). Demand estimation for single family rental housing will need to overcome similar issues faced by Bayer et al. in making apriori/empirically based assumptions of different "product" types and submarket geographies.

In this paper I interpret the heterogeneity in post-merger price effects to reflect differences in marginal cost efficiencies, but only back up this claim with weak theoretical and empirical support. Following structural model estimation, I would be able to back out marginal costs by fitting the model's predicted post-merger outcomes to the realized price outcomes of the mergers I study in this paper. Another way to test this mechanism is to estimate equation (16) on per-unit operating costs. Other possible mechanisms would needed to be tested and ruled out as well, such as product repositioning by the merged firm and differences in consumer preferences across local markets.

6.3 Conclusion

Reporting around the recent rise in mergers among large players in the single family rentals (SFR) space has mostly emphasized the increase in market power that these merging firms would be able to exploit to raise rents. As the extant literature on mergers and my estimates show, however, the reality is a bit more complicated. Though mergers do induce upwards pressure on price due to the merged firm internalizing the competition between its constituent firms, this upwards pressure can be offset or even reversed if the merger also creates cost efficiencies. With SFR mergers, I show that the low levels of concentration and pre-merger market shares result in very modest and nearly economically insignificant changes in post-merger rent prices, with estimated effects in even the most extreme cases of a rent increase of under 10% in the case of the Starwood–Invitation merger, and a rent decrease of 20% for the former Colony American Homes properties in the Colony–Starwood merger. Perhaps more interesting is the heterogeneity in price effect that I estimate for each of the mergers. I demonstrate that this results in predictions of significant merger-induced cost efficiencies for the target properties in the Colony–Starwood merger, and significant merger-induced operating cost increases for the target properties in the Starwood–Invitation merger, using the results from my 3-firm model.

My conclusions are subject to a number of caveats as result of my sample construction and identification strategy. I assume local SFR markets compete within zip codes, but this could actually be overestimating the geographic area of local competition, and thus underestimating the extent of concentration in these markets. Any price changes that result from the mergers that are common across all markets and uncorrelated with pre-merger shares would not be reflected by my estimates, though this seems like an unlikely occurrence in my setting, given that rents are known to be closely tied to neighborhood factors, such as local wage levels and demand for the area.

Nevertheless, this paper makes important steps in evaluating the ex-post effects of three large, recently completed SFR mergers, and demonstrates the importance of considering heterogeneous price responses and the influence of merger-induced efficiencies when conducting such evaluation. Though I conclude that the unilateral competitive effects of these past mergers on prices has been virtually negligible, there is still much work that needs to be done in order to determine whether future SFR mergers will be similarly innocuous, and whether only specific properties in a market are at risk of seeing large price hikes in response to changes in competitive structure. Single family rentals continues to be a growing area of investor interest, and it is likely that M&A activity will continue ramping up in this space for the foreseeable future.

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Appendix A Event-Time Trend Plots

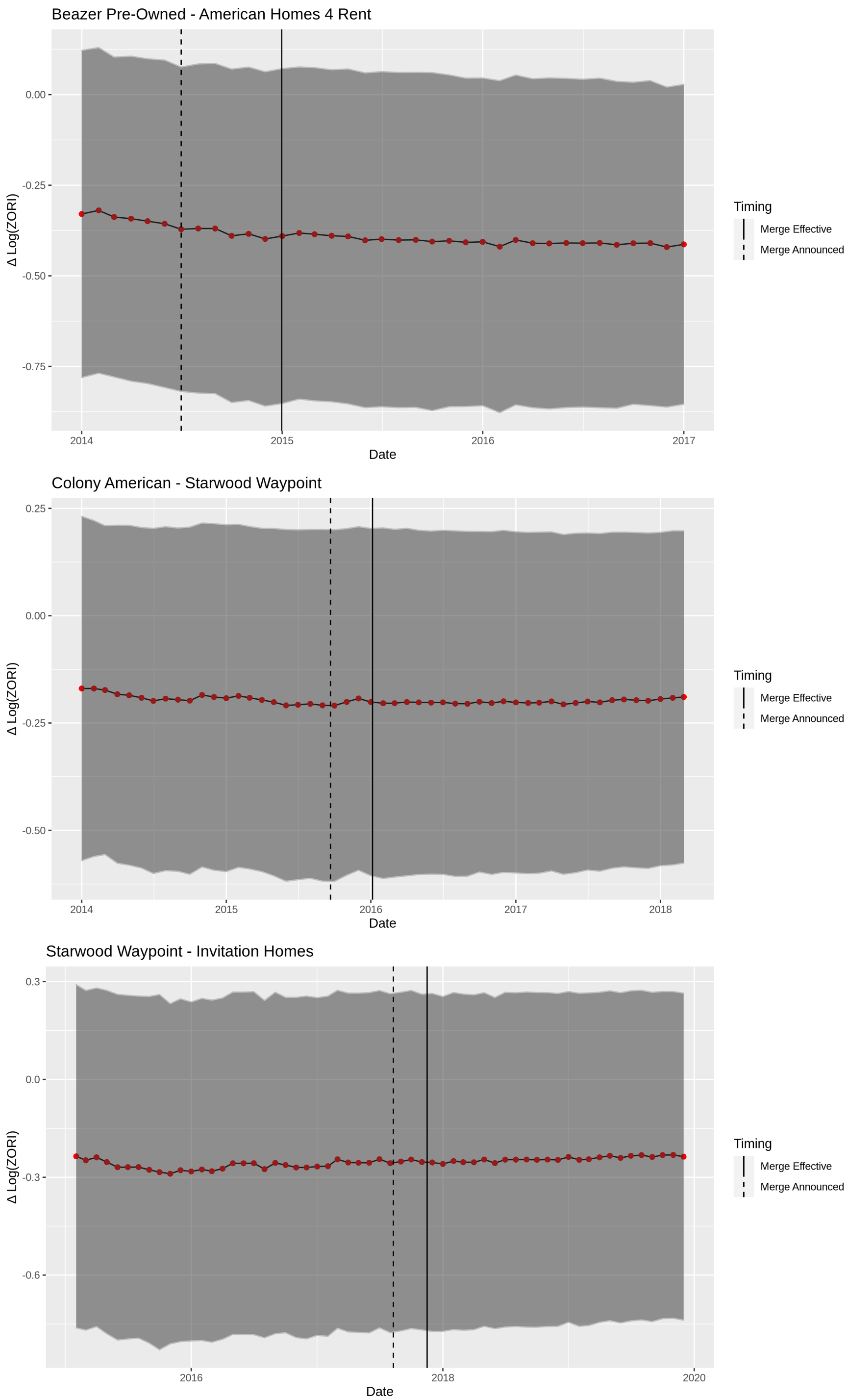

Figure 8: Difference in Log(ZORI) Trends Between Treated and Control Zips

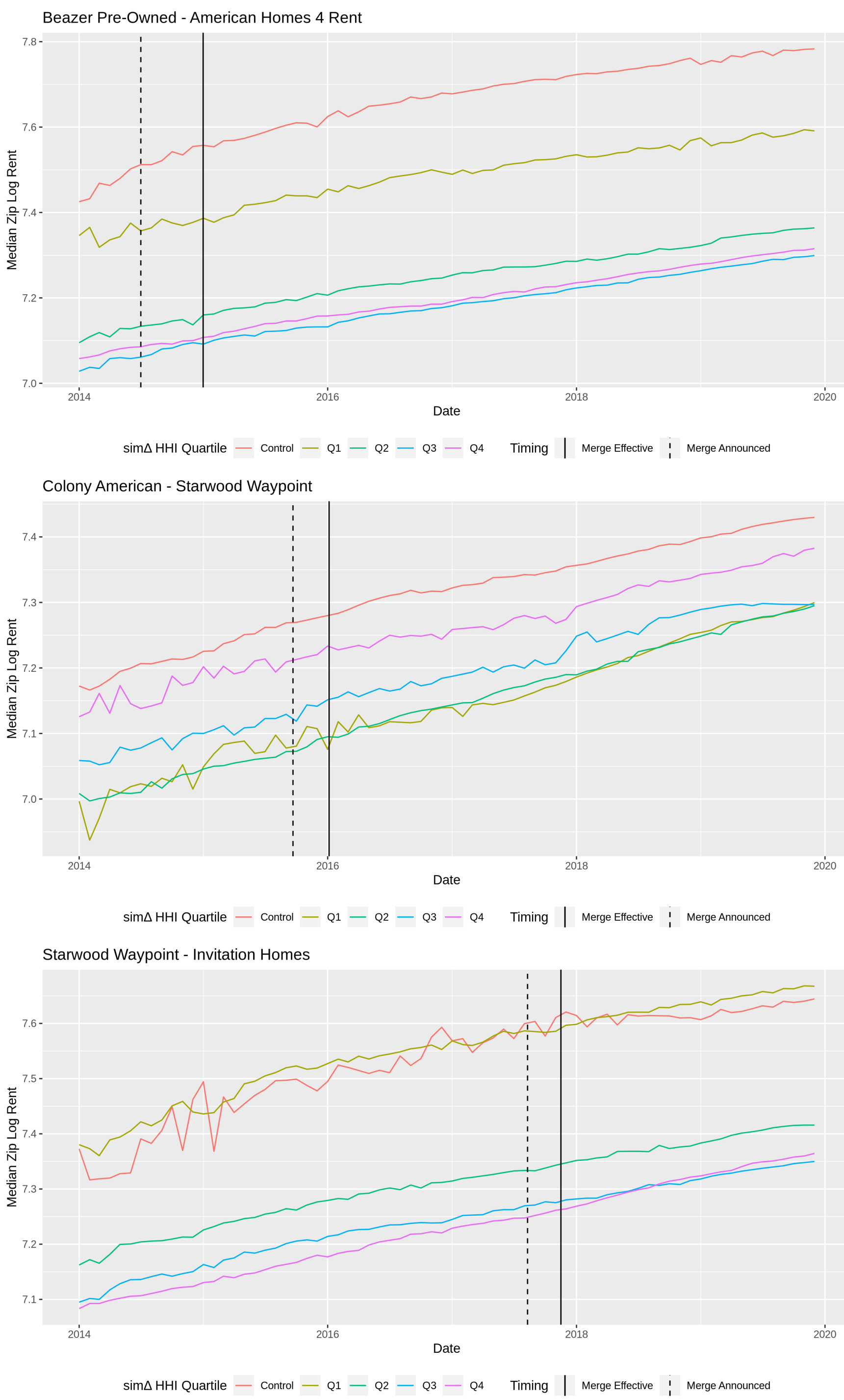

Figure 9: Log(ZORI) by $sim\Delta HHI$ Quartile

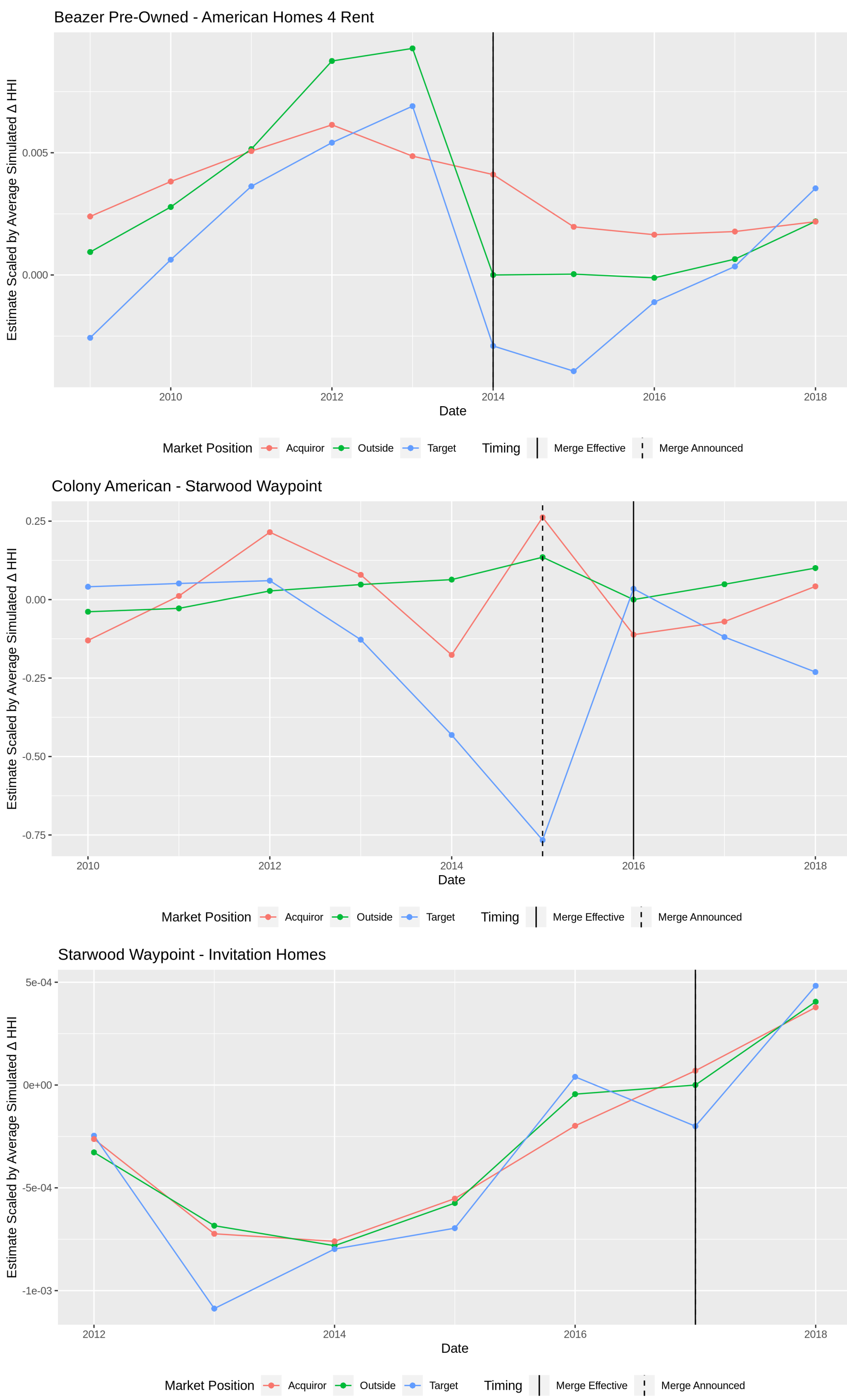

Figure 10: Estimates of (19) on Property Data with Firm Dummies

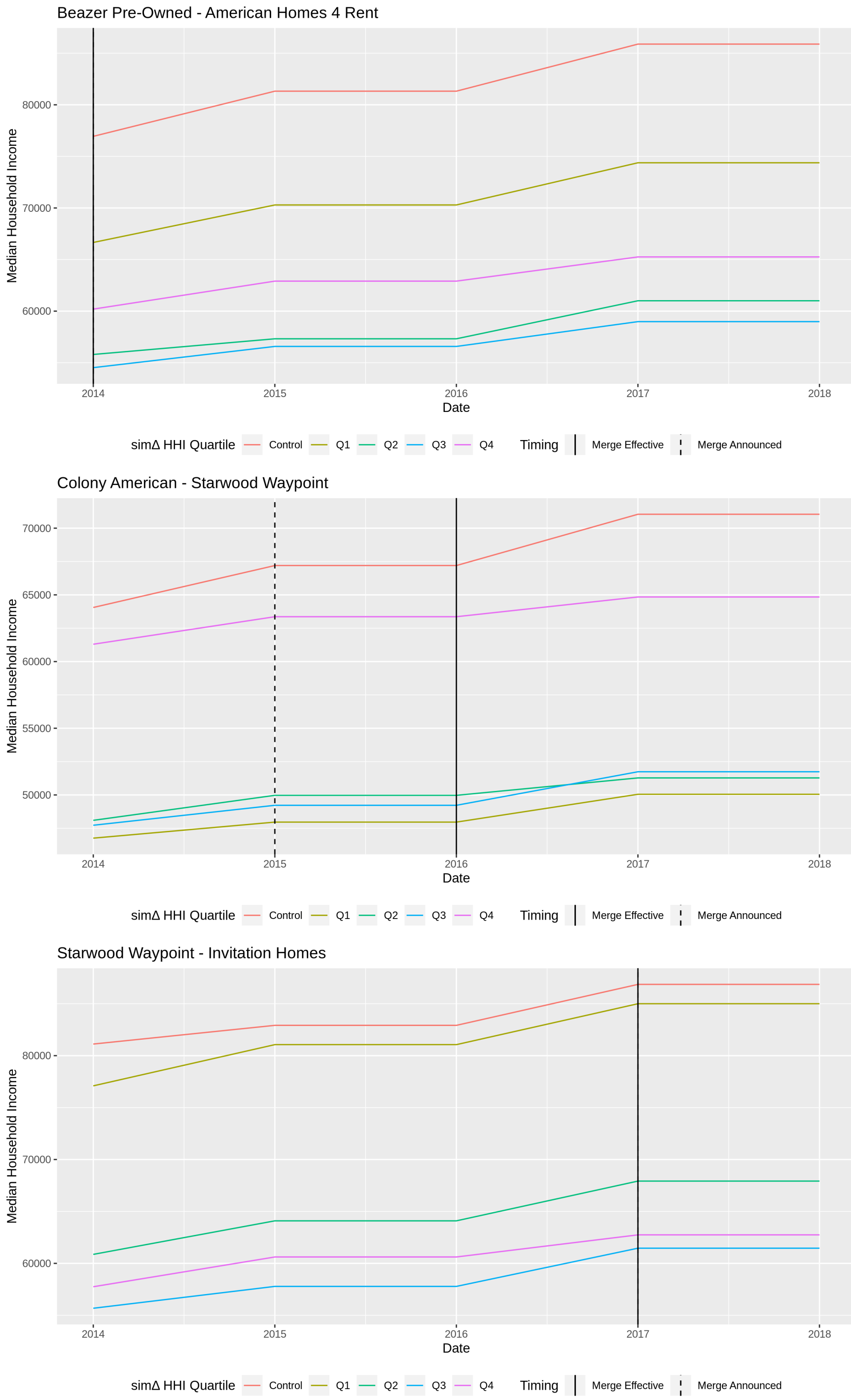

Figure 11: Median Household Income by $sim\Delta HHI$ Quartile

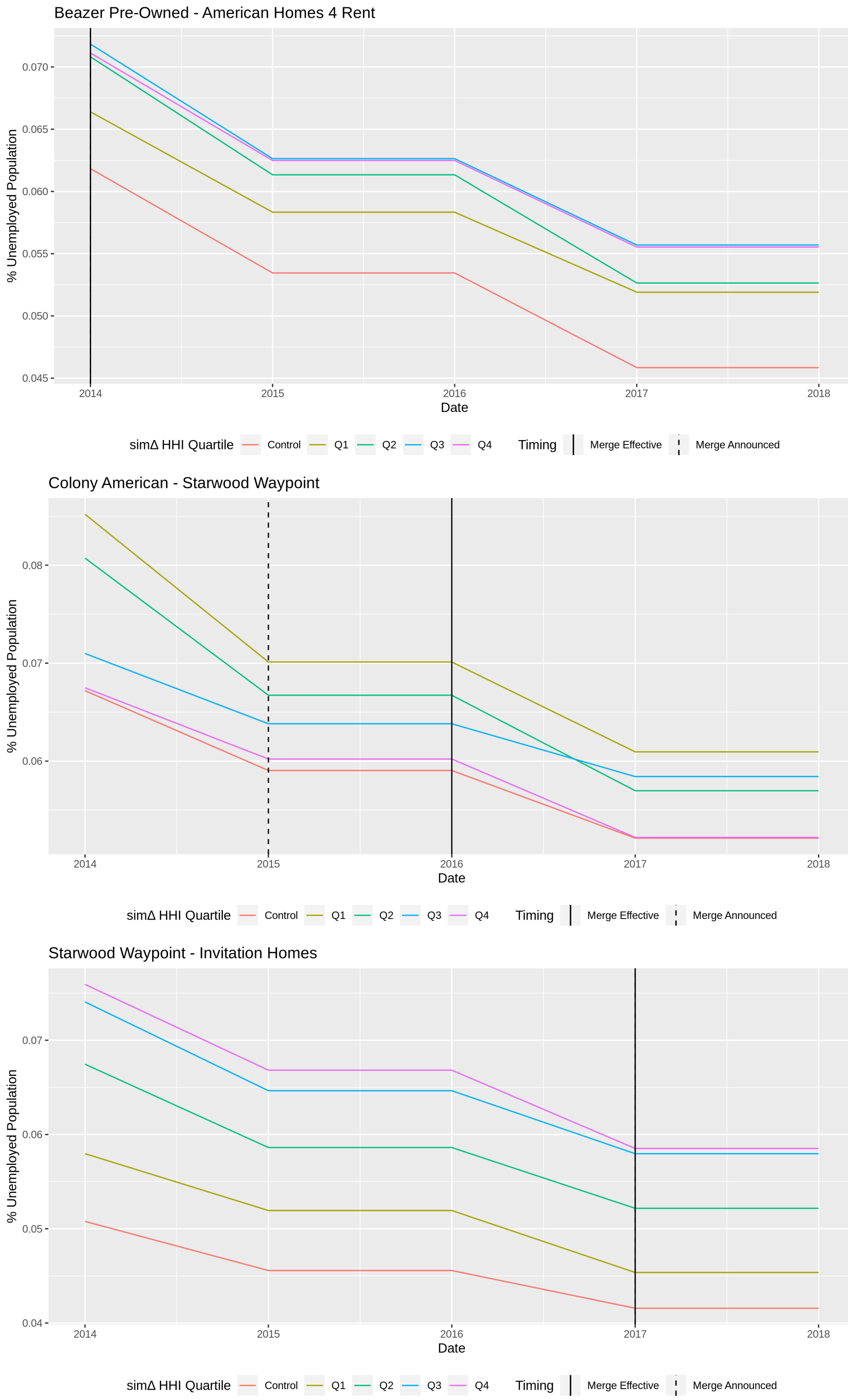

Figure 12: Unemployment by $sim\Delta HHI$ Quartile

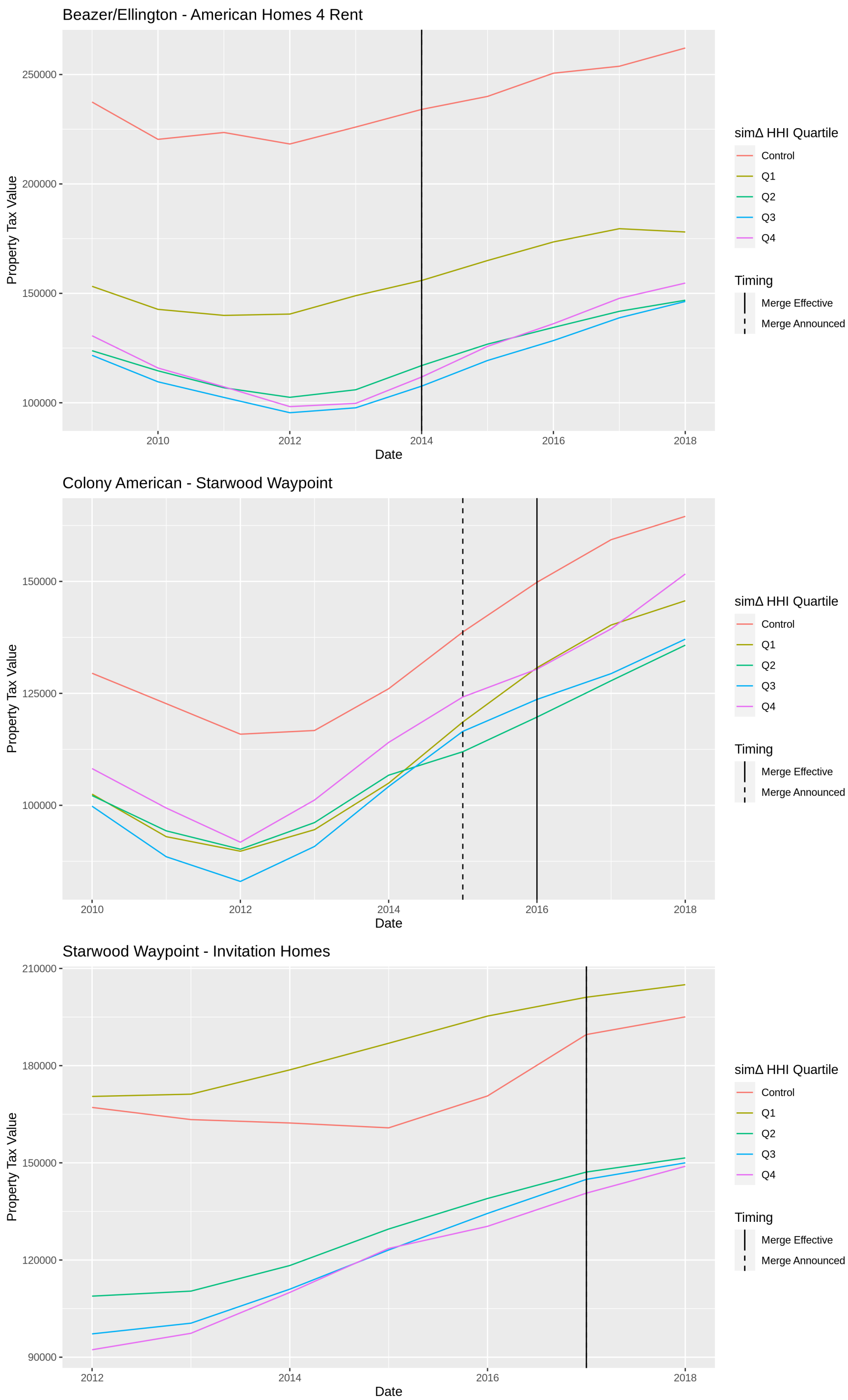

Figure 13: Property Tax Value by $sim\Delta HHI$ Quartile

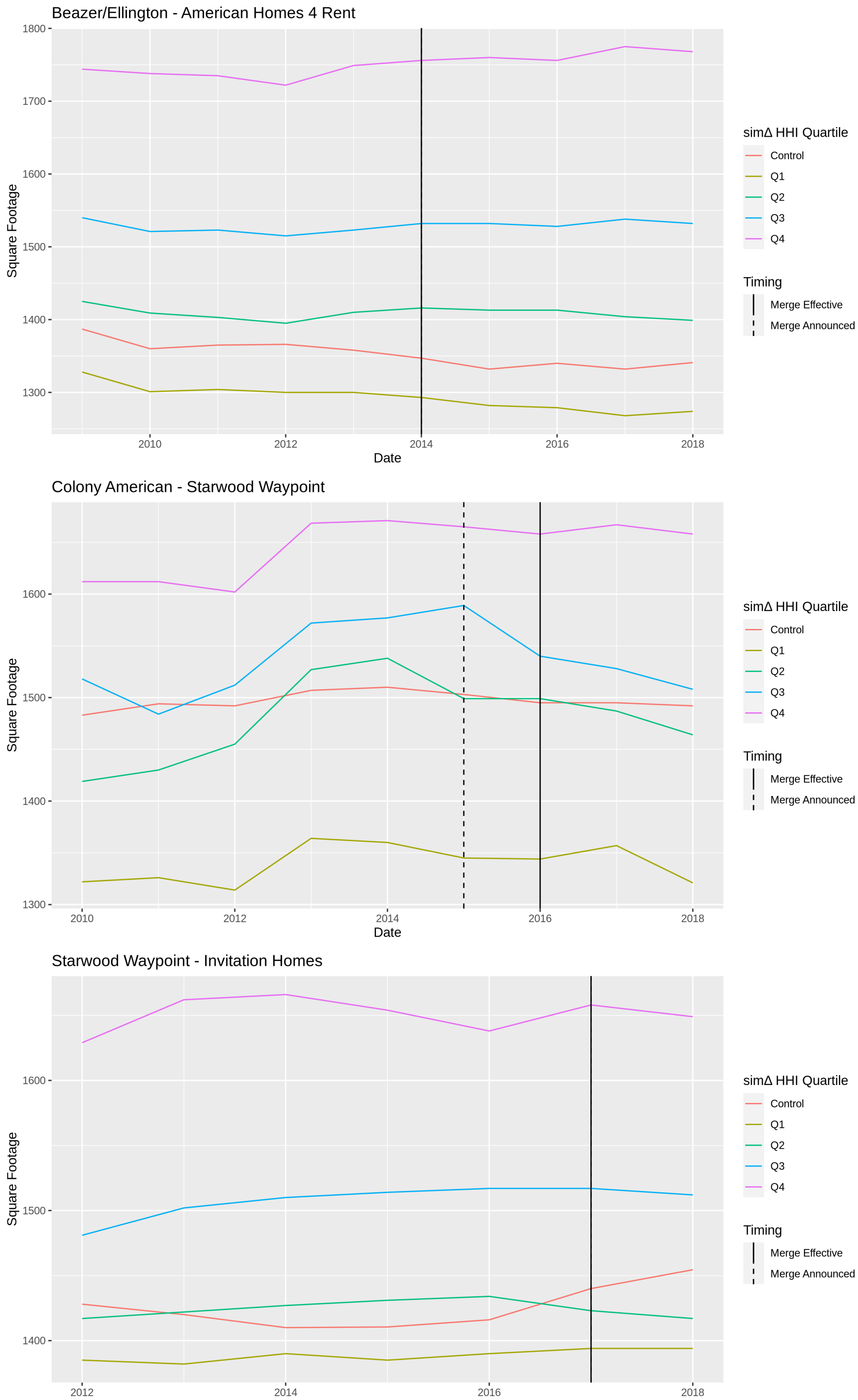

Figure 14: Property Square Feet by $sim\Delta HHI$ Quartile

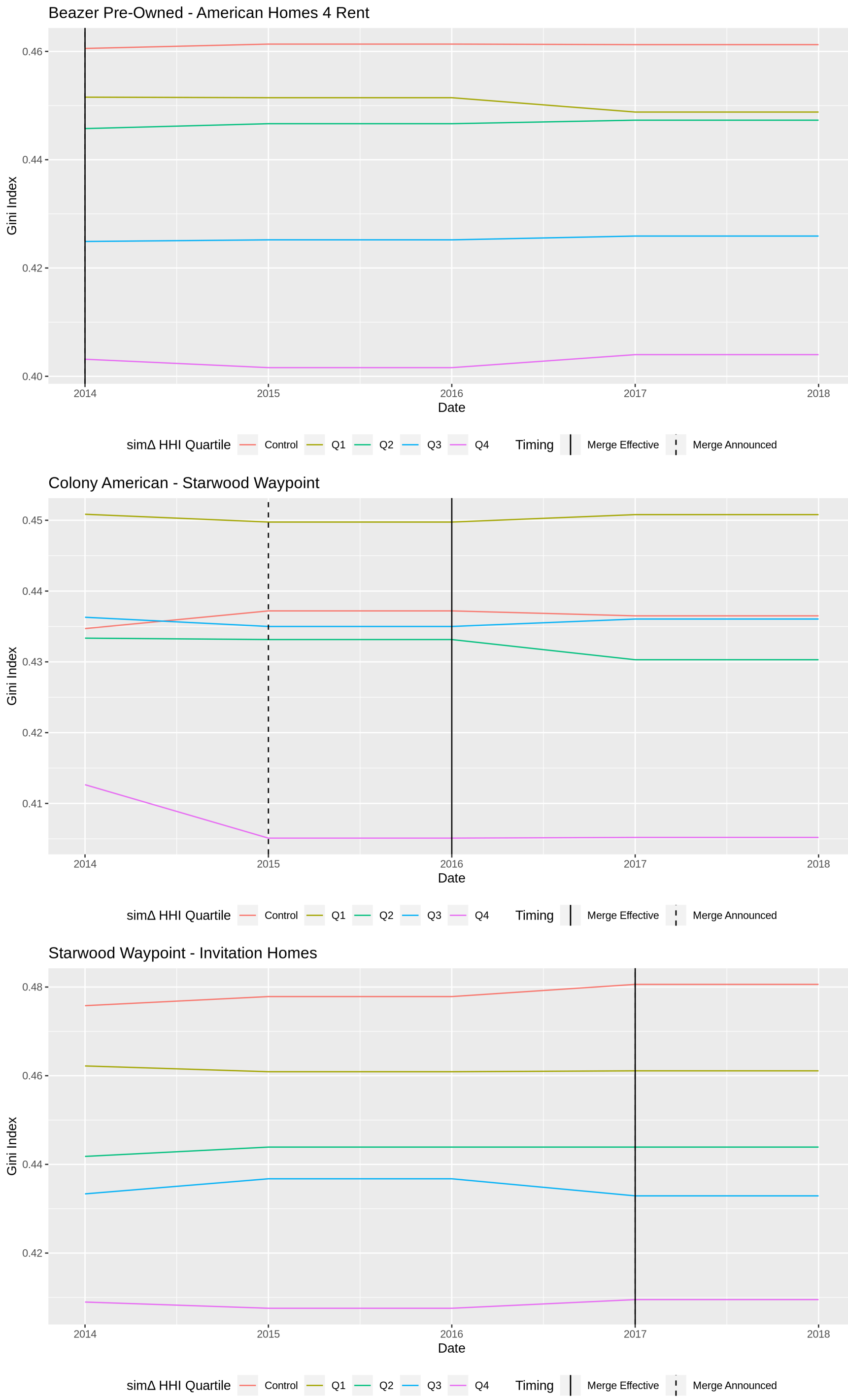

Figure 15: Gini Inequality Index by $sim\Delta HHI$ Quartile

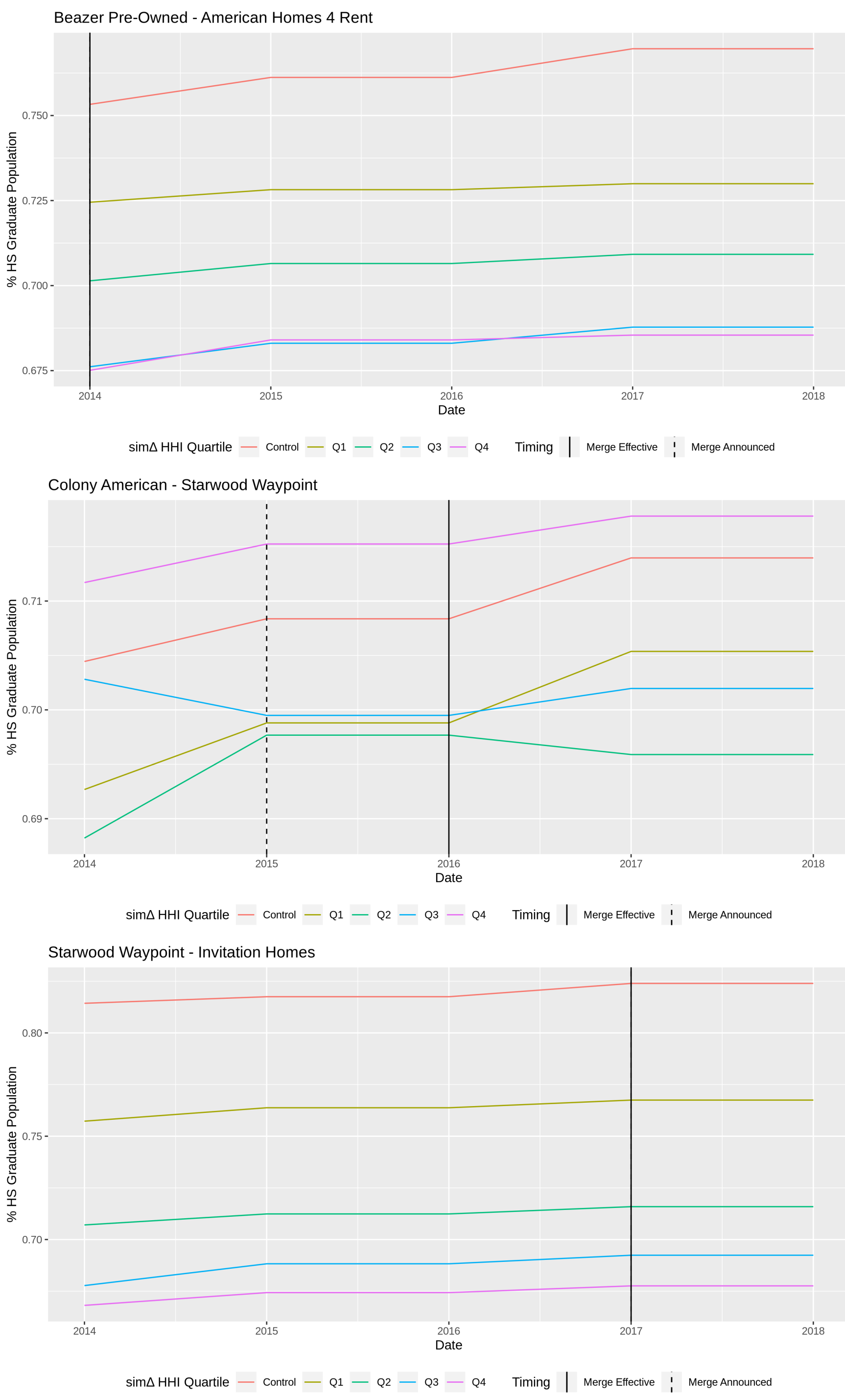

Figure 16: % HS Graduates by $sim\Delta HHI$ Quartile

Appendix B Price Estimates Without $sim\Delta HHI$ Interaction

Table 8: Effect of Mergers on Zip Code Prices

	Coefficient: $sim\Delta HHI * Post$				
	OLS	Property Controls	ACS Controls	Zip Trend	
Beazer – AH4	-0.031^{***}	-0.025^{***}	-0.027^{***}	-0.027^{***}	
	(0.003)	(0.003)	(0.003)	(0.003)	
	[127,338]	[127,338]	[105,408]	[105,408]	[105,408]
Colony – Starwood	-0.016	0.015	-0.012	-0.012	
	(0.013)	(0.013)	(0.010)	(0.010)	
	[128,017]	[128,017]	[105,960]	[105,960]	[105,960]
Starwood – Invitation	0.016^{***}	0.022^{***}	0.018^{***}	0.018^{***}	
	(0.005)	(0.005)	(0.004)	(0.004)	
	[130,748]	[130,748]	[108,205]	[108,204]	[108,205]

Notes: The table reports estimates of (16) excluding $sim\Delta HHI$, for each merger. *, **, *** indicate significance at the 0.1, 0.05, and 0.01 levels respectively. Standard errors for the coefficients are reported in parentheses, and sample sizes are reported in brackets. The average $sim\Delta HHI$ are 0.71, 0.09, and 2.37, for the Beazer–AH4, Colony–Starwood, and Starwood–Invitation mergers, respectively.

Table 9: Effect of Mergers on Property Prices

	Property/Zip Controls			Zip Trend			Polynomial Zip Trend		
	Outside	Target	Acquiror	Outside	Target	Acquiror			
Beazer - AH4	-0.0114^{*}	-0.0188^{***}	-0.0178^{**}	0.0242^{***}	0.0236^{***}	0.0286^{***}	-0.009	-0.015^{**}	-0.0119
	(0.0059)	(0.0065)	(0.0082)	(0.0080)	(0.0085)	(0.0093)	(0.0059)	(0.0065)	(0.0082)
		[2,825,812]			[2,825,810]			[2,825,810]	
Colony - Starwood	0.0336^{***}	-0.0195^{**}	0.0095	0.0368^{***}	-0.0148^{*}	0.0284^{***}	0.0403^{***}	-0.0192^{**}	0.0091
	(0.0043)	(0.009)	(0.0087)	(0.0076)	(0.009)	(0.0095)	(0.0046)	(0.008)	(0.0087)
		[2,604,538]			[2,604,536]				
Starwood - Invitation	0.0577^{***}	0.0663^{***}	0.0355^{***}	-0.0034	-0.0023	0.0070	0.0552^{***}	0.0604^{***}	0.0353^{***}
	(0.0099)	(0.0107)	(0.0107)	(0.0094)	(0.0102)	(0.01)	(0.01)	(0.0107)	(0.0106)
		[2,150,581]			[2,150,577]			[2,150,577]	

Notes: The table reports estimates of the coefficients β_O, β_A, and β_T in (16), excluding $sim\Delta HHI$, for each merger. *, **, *** indicate significance at the 0.1, 0.05, and 0.01 levels respectively. Standard errors for the coefficients are reported in parentheses, and sample sizes are reported in brackets. The average $sim\Delta HHI$ are 0.75, 0.01, and 3.38, for the Beazer–AH4, Colony–Starwood, and Starwood–Invitation mergers, respectively.

Appendix C Derivation of 3-firm compensating marginal cost reduction

The formula for computing the CMCR is

$$\Delta \boldsymbol{c}_{comp} = -\boldsymbol{B}^{-1}\boldsymbol{z}(\boldsymbol{p}^0)|_{\boldsymbol{c}_{pre}} \tag{20}$$

In the 3-firm setting with a merger between firms 1 and 2, we have

$$\boldsymbol{B} = \begin{pmatrix} -\frac{\partial D_1}{\partial p_1} & -\frac{\partial D_2}{\partial p_1} & 0 \\ -\frac{\partial D_1}{\partial p_2} & -\frac{\partial D_2}{\partial p_2} & 0 \\ 0 & 0 & \frac{\partial D_3}{\partial p_3} \end{pmatrix}$$

The first order derivatives $\boldsymbol{z}$, when evaluated at pre-merger prices and marginal costs in the post-merger ownership structure become

$$\begin{aligned}
\frac{\partial \pi_1}{\partial p_1} &= D_1 + (p_1 - c_1')\frac{\partial D_1}{\partial p_1} + (p_2 - c_2')\frac{\partial D_2}{\partial p_1} = (p_2 - c_2')\frac{\partial D_2}{\partial p_1} \\
\frac{\partial \pi_2}{\partial p_2} &= D_2 + (p_2 - c_2')\frac{D_2}{\partial p_2} + (p_1 - c_1')\frac{\partial D_1}{\partial p_2} = (p_1 - c_1')\frac{\partial D_1}{\partial p_2} \\
\frac{\partial \pi_3}{\partial p_3} &= D_3 + (p_3 - c_3')\frac{D_3}{\partial p_3}
\end{aligned}$$

as a result of the pre-merger equilibrium condition. We therefore have

$$-\boldsymbol{B}^{-1}\boldsymbol{z}(\boldsymbol{p}^0)|_{\boldsymbol{c}_{pre}} = \frac{1}{\frac{\partial D_3}{\partial p_3}\left(\frac{\partial D_2}{\partial p_2}\frac{\partial D_1}{\partial p_1} - \frac{\partial D_1}{\partial p_1}\frac{\partial D_2}{\partial p_1}\right)} \begin{pmatrix} (p_2 - c_2')\frac{\partial D_2}{\partial p_1}\frac{\partial D_2}{\partial p_2}\frac{\partial D_3}{\partial p_3} - (p_1 - c_1')\frac{\partial D_1}{\partial p_2}\frac{\partial D_2}{\partial p_1}\frac{\partial D_3}{\partial p_3} \\ (p_1 - c_1')\frac{\partial D_1}{\partial p_2}\frac{\partial D_1}{\partial p_1}\frac{\partial D_3}{\partial p_3} - (p_2 - c_2')\frac{\partial D_2}{\partial p_1}\frac{\partial D_1}{\partial p_2}\frac{\partial D_3}{\partial p_3} \\ 0 \end{pmatrix}$$

Substituting in elasticities for the first derivatives and canceling out terms brings us to the derived result

$$\Delta \boldsymbol{c}_{comp} = \frac{1}{\epsilon_{22}\epsilon_{11} - \epsilon_{21}\epsilon_{12}} \begin{pmatrix} \epsilon_{21}(\frac{D_2}{D_1}m_2 p_2 \epsilon_{22} - m_1 p_1 \epsilon_{12}) \\ \epsilon_{12}(\frac{D_1}{D_2}m_1 p_1 \epsilon_{11} - m_2 p_1 \epsilon_{21}) \\ 0 \end{pmatrix}$$

Appendix D Sample Construction: Imputation Procedure

Though the raw Corelogic tax data is technically supposed to a full panel of records, data is often spotty for some states for arbitrary years. Furthermore, even for properties for which we observe a full set of tax years, the values are often missing for key variables. Thus, when I construct my full sample by linking the tax and deed records, I also expand the sample panel to a fixed range of years from 2000-2018 and impute the data fields for each property over this period. I run two different procedures depending on the type of variable I impute. For ownership variables (e.g. owner name, owner address) for each property, I impute the values as follows:

1. Forward-fill all non-missing values up until either the next observation of a non-missing value or the next observation of a new deed transaction
2. Similarly backwards-fill the non-missing values until either the first year of a deed transaction or the property built year is encountered

For property characteristic variables (e.g. square feet, tax value), I impute ignoring the boundaries between deed transactions:

1. Forward-fill all non-missing values up until the next observation of a non-missing value
2. Backwards-fill the first non-missing value until the cutoff year (2000) or the property built year

Appendix E SFR Properties in All States in Sample