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Do house prices impact business starts? ☆

Lakshmi Balasubramanyan [a,*], Edward Coulson [b]

[a] The Federal Reserve Bank of Cleveland, Indiana State University, United States
[b] The Pennsylvania State University, United States

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ABSTRACT

At the national level, business starts and housing prices both fell dramatically over the 2007–2009 period. Using a proprietary database of business starts this paper quantitatively models the interaction between house price and business starts from 2005 to 2009. We identify the impact by exploiting the cross-sectional variation in house price changes during the period. Controlling for observable and unobservable city characteristics, we find the significance of a robust relationship between house prices and business starts depends on the size of the business starts; a robust link exists between house prices and very small business, whereas, no significant robust link is seen for large business starts.

1. Introduction

Business starts declined dramatically over the 2007–9 period. Such a drop is not atypical in recessions—Tuna (2009) notes that starts declined by about 9% in the recession of 2001—but the sheer scale of the decline in 2007–09 is unprecedented. The proprietary Dunn and Bradstreet Metropolitan Statistical Area (MSA) business starts data suggest that over the two years in question, business starts declined by over 60% (Fig. 1). The coincident decline of the housing market over the same period suggests a link between housing prices and business starts. Such a link may arise through the use of real estate, and more particularly, owner-occupied housing as collateral.

Substantial anecdotal evidence for this phenomenon exists. In an online commentary, Shane (2010) points out that "falling real estate prices impinge on the ability of small employers to borrow the money they need to fund their operations because small businesses use real estate to obtain credit in a variety of ways." Dennis (2010) notes that 95% of small business owners own real estate (either residential or commercial or investment properties) which can serve as collateral, and 20% hold mortgages that finance non-real estate business capital. The fall in real estate prices, and particularly the steep drop in residential prices, can therefore have binding effects on business capital formation.

The theoretical link between credit markets and collateral value is established by Bernanke et al. (1999), following Bernanke and Gertler (1989). In that model, the external finance premium depends explicitly on the collateral held by the borrower, since a large amount of collateral reduces the agency cost faced by the lender, and interest rates are lower for highly collateralized borrowers. A decline in the value of the collateral will correspondingly raise rates and lower the number of entrepreneurs. While Bernanke et al. (1999) did not explicitly consider real estate as a source of collateral, such a link is explored by

☆ Views expressed in this paper are not intended to express those held by the Board of Governors of the Federal Reserve System or of any officials of the Federal Reserve Bank of Cleveland.

* Corresponding author.
E-mail addresses: Lakshmi.Balasubramanyan@clev.frb.org (L. Balasubramanyan), fyj@psu.edu (E. Coulson).

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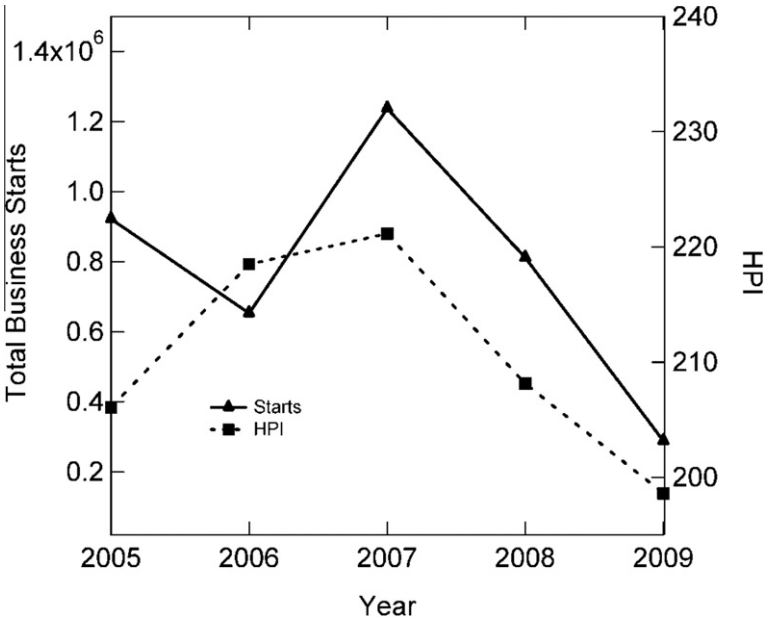

Fig. 1. Plot of house price index (HPI) and total business starts from 2005 to 2009 at the national level. The HPI and total business starts have been aggregated and averaged across the metropolitan statistical areas for 2005–2009.

Chen and Leung (2008) and Jin et al. (2012). In the latter, this mechanism is explicit. A rise in real estate prices raises potential entrepreneurs' collateral, reduces the external finance premium, and in their model this leads to both a reduction in bankruptcy and an increase in the number of entrepreneurs.

The empirical evidence for this link between collateralizable wealth and entrepreneurial behavior is a little mixed, although the balance of the evidence seems to favor the link between house prices and startups. Hvide and Moen (2010), using Norwegian data, find that personal wealth of the entrepreneur influences the size of the startup. Analyzing United Kingdom data, Black et al. (1996) find that an increase in housing equity yields an increase in business starts. Robson (1996) casts doubt on this finding, but demonstrates an empirical link between increasing house prices and declining business failures. As the value of collateralizable property declines, Goodhart (2008) find that firms have difficulty in borrowing to finance business investments. Gan (2007a,b) finds that the land market collapse in Japan provided a shock to collateral value, and had a significant statistical and economic impact on corporate investment via the collateral channel. However, Lusardi and Hurst (2004) find little relationship between housing wealth changes and business starts in the US using micro data from the Panel Study of Income Dynamics.[1]

There is a second, perhaps more direct, channel for housing prices to affect business starts, and that is through cash-out mortgage refinancing. During times of rising housing prices, a common financial strategy was to liquefy housing assets by paying off an existing mortgage, and borrow the full value of the house, pocketing the difference. While research on the uses of these funds has concentrated on the effect of this wealth on consumption expenditure (Bostic and Gabriel, 2009; Gan, 2010) there is evidence (Greenspan and Kennedy, 2008; Canner et al., 2002) that while the majority of these funds were used for consumption (prominently, home improvement) some part of them were for financial investments of various kinds, presumably including new businesses.

Therefore in our empirical model we also include HMDA data on the refusal rates of mortgage loan applications, and the extent to which these refusals were for reasons of insufficient collateral. If the latter has an impact on business starts separate from that of housing prices alone, this is evidence the refinancing channel has shut down along with the collateral channel.

Like several of the empirical papers discussed above, we exploit regional variation in housing price changes to identify its impact on business starts. The recent housing market crash provides especially strong variation, and thus provides an excellent setting for testing the links between housing prices and business starts. However the severity and breadth of the downturn provides several possible causal factors for the decline in business starts, and to the extent that these are correlated with cross-city variation in housing prices, we must provide corresponding controls. We also attempt to distinguish between the role of housing price declines on starts of different size categories. Using standard fixed effects models as well as Tobit

[1] Aoki et al. (2002), Lustig and Nieuwerburgh (2005), and Ortalo and Rady (2004) provide links between house price appreciation and increased consumption through the collateral channel. These studies argue that house price increases can fuel consumption by reducing borrowing constraints. However, it is unclear if the same condition can also fuel business formation.

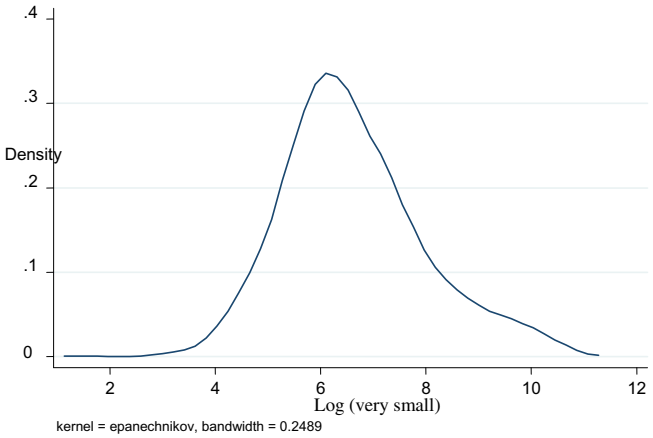

Fig. 2. Density of (log) very small business starts.

and Poisson regressions of starts on various macroeconomic and credit market variables, the results indicate that house prices do indeed impact business starts. Though the effect is small relative to the huge declines in starts at the end of our sample period, very small business starts in particular, those that have between zero and twenty employees, show statistically significant response to changes in house price. These are the starts that one would expect to be most affected by declines in the value of housing collateral. The cash-out refinance channel also appears to be disrupted.

2. Data

The MSA-level business starts data are proprietary data from Dunn and Bradstreet (D&B). The sample is composed of counts of business starts from 2005 to 2009 from 376 MSAs. A business is considered a "start" from the point in time it is registered by the appropriate state agency and receives a business tax identification number.[2] The D&B counts are further categorized based on the number of employees at the time of the startup. Business starts comprised of zero to twenty employees are considered very small business starts. Businesses that have a number of employees between twenty and two hundred employees are considered small businesses, while business starts that have over two hundred to five hundred employees are considered medium businesses and over 500 are large. The vast majority of these starts are in the "very small" category. Fig. 2 displays the smoothed density of the log of the number of very small starts over all year-MSA observations, and it can be seen that this variable is bell-shaped and well approximated by the normal density. The remaining categories are more problematic in that there are substantial numbers of zeroes. The histograms in Figs. 3–5 display this. In Fig. 3 it can be seen that the distribution of small business starts can range up to 800, but the histogram shows a large spike at 0; 12% of the city-year observations have zero small business starts. In Figs. 4 and 5 it can be seen that the "zero problem" is exacerbated for the larger size categories. 81% of the observations had zero medium business starts and 89% had zero large starts. We discuss our response to this feature of the data in the estimation section which follows.

Since we are interested in the loss of collateral due to falling house prices, our measures of prices should be nominal; we need the dollar change in house prices for each city year observation. Nominal price series do not exist for a large number of MSAs in our data. The all-transactions house price index, available from the Federal Housing Finance Administration, is available for each metropolitan area, but this is anchored to a value of 100 for all cities in 1995, and thus does not provide information on nominal losses (or gains) during our time period. Therefore we use median (nominal) house values for the year 2000 from the US Census as an anchor and convert the FHFA index to nominal dollars.[3] Given that we basically have every US metropolitan area in our data set, the all-transactions data is the only one that will serve our purpose—the sales-only index will not provide suitable indexes for all times and locations. This data is available on a quarterly basis; since we are constructing an annual panel, we use the quarter

[2] This definition of a business start is slightly different from that of the Census Bureau. Doms (2011), for example, uses entry into the Longitudinal Business Database as the definition of a start, which requires the establishment to have at least one employee (Bureau of the Census, Center for Economic Studies, undated). Dun and Bradstreet aggregate counts are therefore higher than in Doms (2011). Since we are concerned with the possibility that the initial establishment of a firm may be inhibited by a lack of capital due to falling house prices, and not whether that firm is sufficiently established to hire workers, the Dun and Bradstreet definition seems preferable. (Also note that the publicly available (aggregated) version of the LBD (Business Dynamics Statistics) aggregates to the state level, which is does not correspond to the notion of housing markets which we investigate here.).

[3] We use the formula (2000 Census value) $*$ index value for year t/index value for year 2000. Note further that this does not provide a constant-quality index across cities. The median house is different in different places. Again, this is purposeful, since we want to know the changes in collateral available to the typical entrepreneur in each place. Of course the typical entrepreneur may not own the median house, but we view our measure as an acceptable approximation.

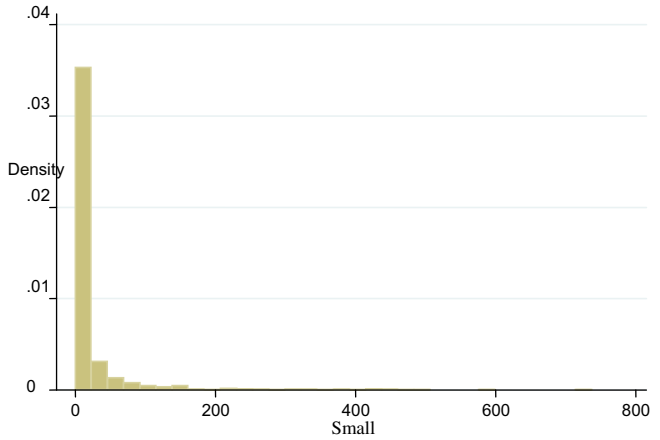

Fig. 3. Distribution of small business starts.

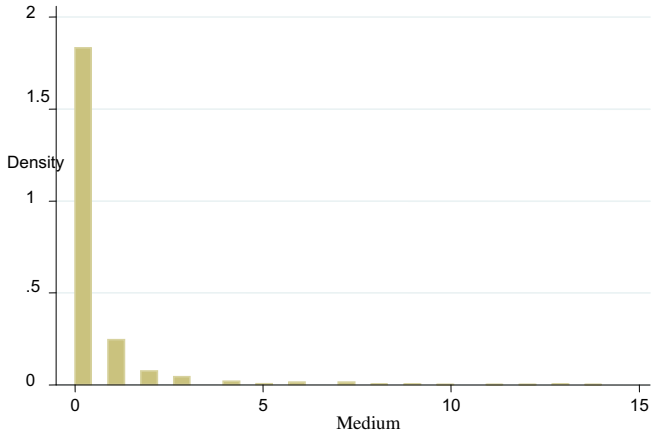

Fig. 4. Distribution of medium business starts.

three index (Q3) for each of the years 2005–2009.[4] Since we wish to estimate the causal impact of house prices on business starts, we must control for other possible channels through which the correlation between these two variables might take place. One obvious channel is local business conditions. Whether or not housing market changes were the source of the downturn, both business starts and housing prices will react to local economic conditions. One natural first step is using MSA fixed effects to control for any time-invariant differences in cities. We also gather data on local unemployment rate from the Bureau of Labor Statistics (BLS). To be consistent with the FHFA, we use Q3 unemployment rates.[5] We also include MSA level gross domestic product in the specifications, obtained from the Regional Economic Accounts of the Bureau of Economic Analysis. We also use population data from 2005 to 2009, obtained from the U.S. Census Bureau, in the model.[6] This is used to put both business starts (where applicable) and metropolitan GDP in per capita terms.

Importantly, we need to control for other sources of variation in credit market conditions across time and across MSAs. We use year fixed effects to model macroeconomic conditions that lead to national changes in credit conditions. However the modeling differences in local credit markets is particularly difficult. The FFIEC's Home Mortgage Disclosure Act (HMDA) Data is potentially helpful in this way.[7] HMDA requires lending institutions to make annual disclosures of their home mortgage and home improvement lending activity, and summaries of this data by MSA and year are available. While the data pertains to the mortgage market, it may provide insight into the creditworthiness of borrowers in the commercial market. We

[4] The choice of third quarter is arbitrary.

[5] The unemployment rate data is obtained from http://www.bls.gov/data/#unemployment.

[6] Population data is obtained from http://www.census.gov/popest/metro/CBSA-est2009-annual.html.

[7] HMDA data is obtained from http://www.ffiec.gov/hmdaadwebreport/.

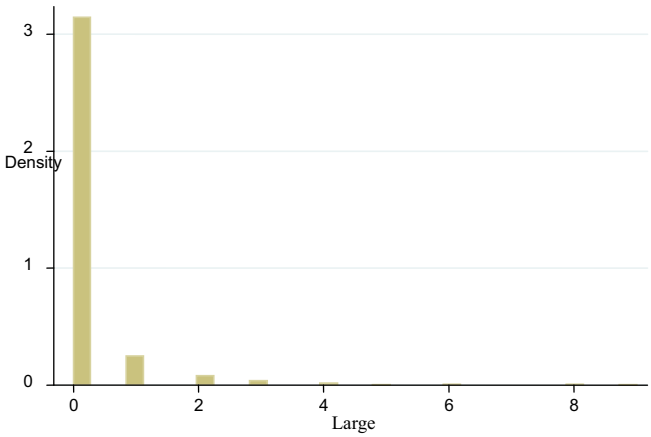

Fig. 5. Distribution of large business starts.

consider the conventional mortgage loan market to be one in disequilibrium, and for which rates are below the equilibrium, as in, say, the credit-rationing models of Stiglitz and Weiss (1981). Gorton (2010) provides insight for this interpretation of market conditions. In such a case, the number of loans approved results from shifts in the supply of loans curve, which in turn arise from (cross-city) shifts in the credit-worthiness of borrowers. Thus we include the number of loans (per 1000 population) for each city-year observation in our regression model. The HMDA summaries also include the number of denials, and moreover a decomposition of this number based on the reason for denial. Among these reasons is "insufficient collateral" which we take to be related to the fall in housing prices. We aggregate all the other reasons into a variable call "non-collateral denials" and include the percentage of denials that were for these non-collateral reasons in the econometric model .If the coefficient of this variable is negative, it is suggestive that credit conditions other than those related to housing prices were also important in business start fluctuations. If it is positive, then collateral-based denials are related to the fall in business starts which we take as somewhat more direct evidence of the importance of cash-out refinancing (or lack of it) being important for business starts. These two variables are obviously imperfect, because they are not direct measures of the credit-worthiness of commercial borrowers in a particular year and MSA, but more direct measures with sufficient coverage of our data do not appear to be available.

Table 1 presents the MSA means, by years, for the above variables, though we omit very small starts because it basically replicates total starts. Several things are of note. The first is (as suggested by Fig. 1) the tremendous decline of starts in 2009. This is accompanied by the decline in the house price index and the increase in unemployment. Note the fundamentally opposite trends in the earlier part of the sample frame.

3. Estimation and results

In Table 2, we present results for models for total business starts (per 1000). The dependent variable is business starts per 1000 population. Because the causality between housing prices (and other variables) and starts is likely not immediate, we lag all the independent variables (with one minor exception below) by one year.[8] In column 1, we simply present the bivariate regression results. As can be seen, starts are highly correlated with lagged nominal house prices—the *t*-ratio is about 12. However there are of course several reasons to doubt any causal interpretation of this coefficient, some of which were alluded to above.

In column 2, we therefore add two economic condition variables: lagged per capita metropolitan GDP (in thousands of dollars) and the lagged unemployment rate. The coefficient of the house price index drops slightly but is still highly significant; house prices still have a role to play in explaining starts, even after taking into account business cycle effects. The roles of GDP and the unemployment rate are entirely expected. Higher unemployment rates and lower (per capita) GDP cause the rate of business starts to fall.[9]

We next take the important step of adding both metropolitan and year fixed effects into the model. Year fixed effects will account for macroeconomic conditions that exist across the various metropolitan areas in the sample, while metropolitan fixed effects will control for city-specific (albeit time-invariant) factors. Column 3 displays this result. The coefficient on house prices rises, and this seems sensible: within-variation of house prices is more important than between-city variation. It is not the case that business starts in San Diego are higher than Dallas because their housing prices are higher. Rather, the estimates suggest that in any given city, declines in starts were due to previous declines in house prices. Equally sensibly, the coefficients on gross metropolitan product and the unemployment rate both decline substantially in magnitude (especially the unemployment rate) and have become statistically insignificant. This suggests that it was the between-city variation in the business cycle effects that was

[8] The results are in large measure not altered using contemporaneous measures, although such regressions may have simultaneity issues.

[9] In point of fact, there is some speculation in the literature that self-employment (which is one form of business startup) is countercyclical (e.g. Becker, 1994; Evans and Leighton (1989) and Dennis (1996)). But we do not find that in these results.

Table 1
This table provides the descriptive statistics at the MSA level the variables used in the regression models.

	2005		2006		2007		2008		2009	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Total starts	2686.50	5155.07	1922.85	4166.60	3678.14	8189.27	2459.24	5435.55	860.56	1986.74
Small starts	49.09	102.59	29.31	65.15	29.15	63.59	15.16	33.97	3.31	8.58
Medium	0.84	2.25	0.61	1.53	0.54	1.54	0.35	1.07	0.14	0.49
Large	0.40	1.24	0.24	0.78	0.22	0.65	0.13	0.44	0.08	0.39
Nominal house prices ($000)	113.46	50.61	120.71	53.74	121.63	51.31	115.64	45.23	111.60	41.56
Unemployment rate	5.16	1.56	4.69	1.49	4.64	1.50	5.76	1.90	9.13	2.78
Per capita GDP	37397.65	10704.14	37899.18	11055.40	38096.32	11215.46	37597.18	11159.66	36450.43	11314.91
Loans per thousand	22.15	13.51	19.81	7.97	14.18	4.35	7.42	2.43	5.05	2.00
Percent of denials due to non-collateral	0.17	0.06	0.18	0.05	0.17	0.06	0.16	0.06	0.14	0.05

important in the previous column—that Detroit had fewer starts than Dallas because Detroit was a perpetually low-income, high-unemployment area, and Dallas was not. The *additional* economic troubles that adhered to Detroit during the crash were not significant in reducing the number of starts, but (to repeat) housing price changes evidently were.

In the next column we add the two credit market variables. The first (lagged total loans per thousand population) is tiny and statistically insignificant. The second (lagged percentage of denials due to non-collateral) is positive and statistically significant. Since we control for total credit quality with the previous variable, the implication is that business starts are lower when collateral (rather than noncollateral) denials are more prominent. This is of interest because collateral denials in HMDA are, naturally enough, largely related to property values (Avery et al., 2012) and suggestive of the importance of the cash-out refinancing as a path to starting a business (Canner et al., 2002) The housing price coefficient does not change very much with the inclusion of these variables, suggesting that the conditional correlation between housing prices and the composition of mortgage denials is low (the unconditional correlation coefficient is 0.06). The next two columns break out two subsamples, each consisting of states particularly affected by the housing crisis. The column labeled "Sand" reports the results of estimating the model using just the states of California, Arizona, Nevada and Florida. These states were associated with the housing and mortgage market overreach that was commonly viewed as fundamental to market breakdown in 2007, and had some of the largest falls in housing prices in the US. Nevertheless, the relationship between house prices and business starts weakens considerably. The coefficient is reduced by half, compared to the previous regression, and becomes statistically insignificant. Fig. 6 sheds some light on that; we provide there a summary of the relationship between our two key variables by plotting the four year percentage changes in house prices and business starts. We highlight California cities in that plot, and it can be seen that the cities of California are prominent outliers. San Diego, Santa Barbara, other coastal areas, and interestingly cities which were considered particularly egregious examples of mortgage overreach—Riverside, El Centro, Salinas and the Central Valley had some of the largest house price declines, yet did not see particularly large declines in business starts. The experience of other sand state MSAs similarly did not conform to any strongly recognizable pattern, hence the weak econometric result.[10] The number of loans does seem to have some correlation with starts—stronger than in the sample as a whole. The sign on the non-collateral denials remains positive though the relationship is a bit weaker than before. All in all, the model appears not to fit the sand states particularly well.

The final column in Table 2 is limited to three rust belt states which were also the site of large numbers of foreclosures, Michigan, Indiana and Ohio. The relationship between house prices and starts is much stronger here. This suggests that differences between strong and weak housing markets within these states is starker than those in the sand states or even the sample as a whole. Fig. 7 displays—similarly to Fig. 6, the four year changes in our two key variables. While the obvious outlier of Detroit draws one's attention, even without that, the difference between cities with both big losses in starts and housing prices (mainly in Michigan), and those that are doing somewhat better (though still with negative changes) is pronounced.[11]

With this basic model in place we stratify the business starts into size categories. Table 3 presents these results. In the first column, we present a fixed effects regression of very small business starts. Since very small starts comprise 99% of total starts we expect this regression to look very much like similar ones in the previous table, and in fact it does. There are no real differences between this set of coefficients and those in column 5 of Table 2.

As noted above, when examining the number of small business starts, we need to deal with the fact that 12% of the observations are zero. We use a Tobit estimator to deal

[10] One could surmise that some business starts were in fact caused by house price declines and foreclosures (e.g. storage facilities). This is an intriguing hypothesis, but one which awaits further research. A cursory examination of sectoral establishment data from County Business Patterns did not reveal anything striking in this way.

[11] The state level patterns in these two figures suggests that parsimony might be achieved by using state level data rather than metropolitan areas. We reestimated the regressions while aggregating up to the state level by using the first-named state in the MSA name. The coefficients were basically the same, Also, in our MSA model we included state-time interactions in the specification. These also did not alter our results.

Table 2
This table presents results for models for total business starts (per 1000). The dependent variable is business starts per 1000 population.

Sample	All	All	All	All	Sand	Rust
Size distr.	Total	Total	Total	Total	Total	Total
Method	OLS	OLS	FE	FE	FE	FE
Nominal House prices ($000)	0.012*** (0.001)	0.009*** (0.001)	0.018*** (0.007)	0.016*** (0.006)	0.008 (0.01)	0.033* (0.017)
Unemployment rate		−0.224*** (0.032)	0.095 (0.054)	0.084 (0.058)	0.147 (0.16)	0.072 (0.08)
GDP per capita ($000s)		0.022*** (0.006)	0.015 (0.017)	0.010 (0.019)	0.014 (0.064)	0.004 (0.029
Loans per 1000				0.008 (.006)	0.027 (0.016)	0.0157 (.021)
% denials for non-collateral				4.367*** (1.17)	3.70 (2.62)	1.81 (3.58)
r2	0.084	0.133	0.705	0.708	0.656	0.889
N	1486	1486	1486	1486	225	156

Notes: "All" refers to all states being included in the estimation. Sand refers to California, Arizona, Nevada and Florida; Rust refers to Michigan, Indiana and Ohio. Total refers to the fact that the entire distribution of firm sizes was included. FE implies that both fixed time and MSA effects were included. Standard errors in parentheses.

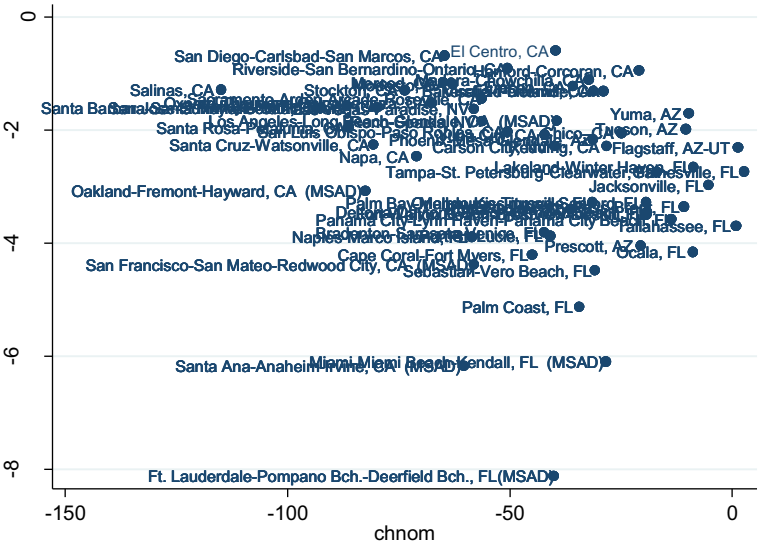

Fig. 6. Change in per capita business starts versus percentage change in house prices, Sand states.

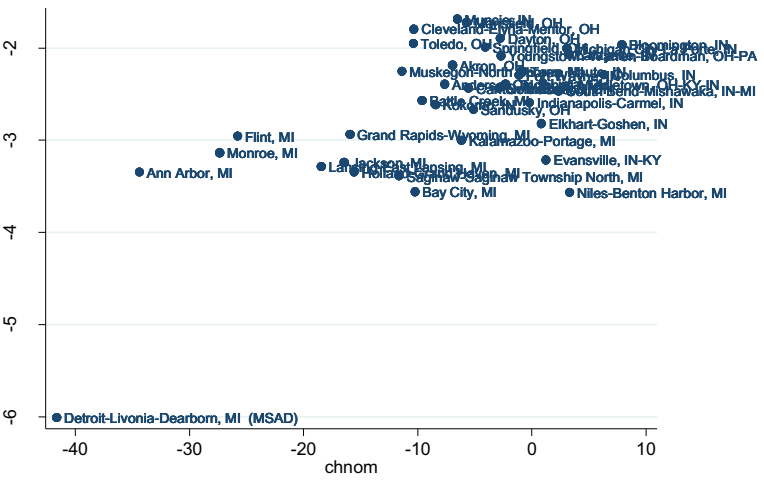

Fig. 7. Change in per capita business starts versus percentage change in house prices, Rust states.

Table 3
Fixed effects, Tobit and Poisson regressions of starts indicated size categories on indicated variables. Fixed time and city effects are included in every specification.

Sample	All	All	All	All
Size distr.	Very small	Small	Medium	Large
Method	FE	FE-Tobit	FE-Poisson	FE-Poisson
Nominal House prices ($000)	0.016*** (.006)	3.77E-05 (5.76E-05)	0.002 (.0056)	−0.0171* (0.008)
Unemployment rate	0.083 (.057)	2.83E-04 (8.02E-04)	−0.102 (0.120)	−0.353 (.214)
GDP per capita	0.011 (.019)	−5.87E-04 (3.29E-04)	−0.068 (.050)	−0.079 (.056)
Loans per 1000	0.008 (.006)	1.22E-05 (1.05E-05	−0.002 (0.007)	−0.25 (0.027)
% denials for non-collateral	4.37*** (1.17)	−0.00140 (0.020)	3.45 (2.41)	−6.28 (3.90)
Population (00,000s)			0.0037 (0.011)	−0.002 (0.02)
N	1486	1486	567	391

with this issue, and the results are presented in column 3 of Table 3. Housing prices are no longer a significant predictor of business starts. The two business cycle variables are insignificant, but the number of loan denials continues to have a positive sign.

The last two columns present the results for medium and large business starts, respectively. Given the distributions of these two variables displayed in Figs. 4 and 5, we use Poisson regressions, which are appropriate for the case when the dependent variable is and integer count of occurrences. Note, though, that in doing so we do not (as in previous models) transform the dependent variable into per capita terms. Instead we use population as an explanatory variable. Note further that we continue to employ both time and city fixed effects. A number of cities had no medium and/or large starts during our sample period; these cities do not, on that account, contribute to the likelihood function and are therefore omitted from their respective regressions.[12] The results are disappointing. In the "medium" regression there is no statistically significant predictor; in particular the coefficient of house prices has the expected positive sign but a *t*-ratio of less than one. There is no real evidence that house prices have the expected impact. In the "large" regression, the coefficient has an unexpected negative sign, although the standard error is somewhat higher than desired for precise inference. None of the other variables seems to have any individual explanatory power.

The story from these regressions is straightforward. Cross-sectional variation in house price changes have a statistically significant correlation with metropolitan business starts. This is the case even after controlling for local business cycles; the residual role for housing prices over and above its role as a signal of local economic conditions is as potential collateral for business startups. Our results suggest, naturally enough, that housing prices exhibit a robust relationship for very small business starts which are the vast majority of starts. While we have no data on the people behind these starts, intuition suggests that these are first-time entrepreneurs whose only source of capital is the home which they occupy, which thereby becomes their source of collateral. Declines in the value of this collateral, according to this story, and the models discussed in the introduction, cause credit to dry up, and the number of starts to decline. Our other credit market measures suggests that this is the major source of cross-city variation in the health of credit markets, especially to the extent that collateral denials are related to housing prices. Larger business starts, on the other hand, seem unaffected by housing price variation. This is presumably due to the fact that for larger enterprises, housing is not the source of collateral; sole proprietors using housing as collateral could scarcely be able to finance a startup larger than a few dozen employees, at best. Even so, none of the other independent variables seemed to have much explanatory power.[13]

Having established that housing prices had an effect on very small business starts, we ask: how much did the fall in housing prices matter? Let us use the extremely large decline in starts from 2008 to 2009 as an example. The unweighted mean decline in housing prices across cities from 2007 to 2008 was approximately $6000. Using the "very small" coefficient from the first column of Table 3 indicates that the fall in starts per thousand population due to the change in housing prices would be roughly $(-6) * (0.0161) = -0.0644$. In a city of 1 million people this would be approximately 97starts which is about six percent of the unweighted average decline in the number of starts, suggesting that the effect is small. While we do not present the coefficient of the time dummies in our table, our inspection of them indicates that they are large contributors to the fit of the model. The coefficient of the 2009 binary is −1.47 which is larger than the mean decline of starts. While house prices at the local level are certainly important, the size of the decline here indicates that general macroeconomic conditions played an even greater role.[14]

[12] This is because the likelihood calculation is conditional on the total number of occurrences, and the likelihood is based on the proportion of occurrences which occur in each year. When that total is zero, these proportions are known (see Wooldridge, 2010). As a specification check, we estimated the mode in per capita terms as a linear regression (with fixed effects) using the full sample and the non-result reported shortly is repeated there.

[13] A referee asked that we provide a robustness check by using the Case Shiller index. We created nominal price series for the 20 cities for which data is available in a manner analogous to our use of FHFA data. The results were much the same. The "all starts" models yielded coefficients on housing prices that were not quite as large in magnitude as Table 3. They were still statistically significant, though the t-stats were not quite as high, given the substantial loss in degrees of freedom.

[14] Even in the regressions dealing with larger-sized startups, the time dummies were large and precisely estimated, and were the primary contributors to goodness of fit.

4. Conclusion

Housing prices fell dramatically over the 2007–2009 period and concurrently business starts declined dramatically indicating a link between housing prices and business starts. By controlling for cross sectional variation, time fixed effects and local macroeconomic conditions; we find that housing prices exhibit a statistically significant relationship with very small and small business starts. Our study lends support to the idea that *de novo* entrepreneurs of very small and small businesses rely on their homes as a source of collateral to access credit. Larger businesses show no response to decline in house prices. However, this is not surprising as large business starts are unlikely to rely on personal housing wealth as a source of collateral or liquidity to acquire credit or funds to fund the business and hence unaffected by variation in housing prices. Though the quantitative impact of house price on business starts is small, our study provides evidence of the collateral channel for small business starts. Shocks to this collateral channel can be a potential feedback mechanism that can affect credit channels and business lending.

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