House Price Gains and U.S. Household Spending from 2002 to 2006

Atif Mian
Princeton University and NBER

Amir Sufi
University of Chicago Booth School of Business and NBER

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Abstract

We examine the effect of rising U.S. house prices on borrowing and spending from 2002 to 2006. There is strong heterogeneity in the marginal propensity to borrow and spend. Households in low income zip codes aggressively liquefy home equity when house prices rise, and they increase spending substantially. In contrast, for the same rise in house prices, households living in high income zip codes are unresponsive, both in their borrowing and spending behavior. The entire effect of housing wealth on spending is through borrowing, and, under certain assumptions, this spending represents 0.8% of GDP in 2004 and 1.3% of GDP in 2005 and 2006. Households that borrow and spend out of housing gains between 2002 and 2006 experience significantly *lower* income and spending growth after 2006.

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Mian: (609) 258-6718, atif@princeton.edu; Sufi: (773) 702 6148, amir.sufi@chicagobooth.edu

Introduction

What was the effect of the large house price gains between 2002 and 2006 on consumer spending? The marginal propensity to consume (MPC) out of housing wealth is either close to zero or very small in benchmark life cycle models.[1] However, an alternative set of *cash-on-hand* models suggest that rising housing wealth is important for spending if it increases access to "cash on hand."[2] These models also predict that the effect of cash-on-hand shocks on spending is strongest for households with low levels of existing cash on hand.

The cash-on-hand view of the housing wealth effect implies that the total spending effect of housing gains depends on how these gains are distributed across the population, and whether homeowners can liquefy the housing gains. This paper estimates the marginal propensity to borrow and the marginal propensity to consume out of housing wealth for a broad spectrum of households, and quantifies the effect on the overall economy. Our results have implications for household behavior at the micro level and business cycle dynamics at the macro level.

We use individual and zip code level data, and exploit cross-sectional variation in house price growth to estimate the impact of rising home values on borrowing and spending. We use the Saiz (2010) housing supply elasticity of a CBSA as an instrument for house price growth, and focus on the heterogeneous effects of house price gains on borrowing and spending.

We sort zip codes by per-capita adjusted gross income in 2002 as our primary measure of cash on hand, but the results are robust to using alternative measures. Our key question is: for the same dollar increase in home value, do low income zip codes borrow and spend more aggressively? Our empirical strategy seeks to estimate this cross-derivative while holding non-housing wealth constant. The housing supply elasticity instrument helps us in this regard since

[1] See Sinai and Souleles (2005) and Campbell and Cocco (2007)
[2] For example, Deaton (1991), Carroll(1992), Carroll and Kimball (1996), Harris and Laibson (2002), and Kaplan and Violante (2014).

contemporaneous wage growth shocks in a CBSA are uncorrelated with its housing supply elasticity.

We find that households in low income zip codes are more likely to borrow out of an increase in home values. A 20% increase in house price growth (about one standard deviation) in a zip code from 2002 to 2006 is associated with a 3 percentage point rise (about two-third standard deviation) in the annual share of outstanding mortgages that are refinanced in a cash-out transaction, where equity is withdrawn from the home. This effect is driven almost entirely by zip codes where the average 2002 income was less than $50 thousand per household. Among zip codes with average income more than $100 thousand, the cash-out refinancing sensitivity to rising house prices is almost zero.

We use individual level data to estimate the marginal propensity to borrow out of housing wealth for homeowners. On average, the marginal propensity to borrow for a homeowner is about $0.19 per dollar of home value increase. But there is strong heterogeneity in this effect: for homeowners with the lowest cash on hand, the marginal propensity to borrow is $0.26. For households with the highest cash on hand, the marginal propensity to borrow is close to zero.

These results are consistent with the cash-on-hand framework; lower cash-on-hand households most aggressively liquefy their home equity when home values rise. Do they also consume? Due to data limitations, we can only answer this question at the zip code level and in the context of new auto purchases. We find strong effects. The average marginal propensity to spend on new autos is $0.02 per dollar of home value increase. This effect is $0.03 for zip codes where households have an average 2002 income of $35 thousand or less. The marginal propensity to spend out of housing wealth shocks is *zero* for households living in zip codes in which the average income is $100 thousand or greater.

We estimate that the *entire* effect of housing wealth on spending from 2002 to 2006 is due to borrowing against home equity. Households spend when home equity rises because it facilitates borrowing. In other words, the "housing wealth effect" is primarily driven by those who are constrained by low levels of cash on hand.

Section 5 of our paper discusses the macro level impact of our cross-sectional estimates. Under some assumptions that we specify in detail, spending against home equity shifts aggregate spending by 0.08% of GDP in 2003, 0.8% of GDP in 2004, and 1.3% of GDP in both 2005 and 2006. We then outline a simple New Keynesian model to help understand the possible *general equilibrium* impact of the shift in aggregate spending. We discuss how the equilibrium outcome depends on the inflationary impact of our estimated shift in aggregate spending and provide evidence in this regard.

Finally, we provide some evidence on the ex-post outcomes of high borrowing and spending against house price gains of 2002 to 2006. In the years after the housing boom, low income zip codes in inelastic housing supply cities see a dramatic relative *reduction* in both income and spending on autos. In other words, households that *ex ante* borrowed and consumed the most *ex post* did not see stronger income growth--in fact, it plummeted. The higher spending on autos from 2002 to 2006 followed by a collapse in spending from 2006 to 2009 implies higher overall spending volatility.

Our paper contributes to the literature on consumer theory and the macroeconomic effects of financial shocks. Our results show that the housing wealth effect on spending is primarily driven by new borrowing and that there is strong heterogeneity in this effect by income and wealth. [3] The results illustrate why the *joint distribution* of wealth and financial shocks matters

[3] The broader literature on housing wealth effect is far too large to be fully summarized here, but it includes Disney, Gathergood, and Henley (2010), Muellbauer and Murphy (1997), Attanasio and Weber (1994), Lehnert (2004),

for the macro economy. A consideration of the joint distribution helps explain some otherwise puzzling facts. For example, the total housing market decline of 2007 to 2009 was similar in magnitude to the crash in equity values in 2001. Yet the macroeconomic effects were *very* different. Our results offer a simple explanation: most of stock market wealth is held by the top-end of the wealth distribution with a very low MPC out of wealth. Similarly, the house price recovery from 2011 onwards did not contribute as much to economic activity as the 2002 to 2006 housing gains. Our results suggest that this might be because the borrowing channel was effectively shut down for those most responsive to house price gains.

Other studies have documented heterogeneity in the marginal propensity to consume out of income and credit availability shocks, such as Parker (1999), Souleles (1999), Gross and Souleles (2002), Johnson, Parker, and Souleles (2006), Parker, Souleles, Johnson, and McClelland (2013), and Jappelli and Pistaferri (2014). The innovation here is to focus on the rise in home values. We believe we are the first to document the strong heterogeneity across the income distribution in the response of spending to an increase in house prices.

Our result on the heterogeneity in marginal propensity to borrow against house price gain is also found in Mian and Sufi (2011). But in that paper we could not estimate the effect on spending, which is critical for understanding the possible real effects of housing boom on the overall economy. Our finding of heterogeneity in marginal propensity to spend is also found in Mian, Rao, and Sufi (2013) for the 2006-2009 period of housing collapse. But the focus in this paper is on cash on hand models when house price increases and on the macroeconomic consequences of housing booms. In this regard, an important contribution of this paper is that it translates the cross-sectional estimate of the marginal propensity to borrow and spend out of

Case, Quigley, and Shiller (2005 and 2013), Haurin and Rosenthal (2006), Campbell and Cocco (2007), Greenspan and Kennedy (2007), Bostic, Gabriel, and Painter (2009), Carroll, Otsuka, and Slacalek (2011), Guerrieri and Iacoviello (2013), and Zhou and Carroll (2012).

housing gains into the possible macro-level impact of house price increases on total spending. The question of translating micro data estimates – that mostly rely on cross-sectional comparisons – to macro aggregates is a very important but difficult question in practice.

The rest of this study proceeds as follows. The next section presents the data and summary statistics. Section 2 presents the theoretical framework that motivates our estimation, and discusses the empirical strategy. Section 3 presents the results on borrowing. Section 4 presents the results on spending. Section 5 quantifies the aggregate impact of house price increase and also discusses the possible general equilibrium consequences of our estimate. Section 6 examines outcomes after the housing boom, and Section 7 concludes.

1. Data and Summary Statistics

Our analysis focuses on quarterly zip code panel of 5,163 U.S. zip codes for which we have data on all key variables from 2000 through 2012. The sample is limited by the availability of house price data from CoreLogic and housing supply elasticity information from Saiz (2010). These two sample restrictions eliminate zip codes in rural areas. More than 90% of zip codes in the sample are classified as urban in the 2000 decennial census, whereas 25% of zip codes not in the sample are urban. Only urban zip codes with a large number of transactions are covered in zip-code level house price indices. These 5,163 zip codes represent just over 50% of the total population of the United States, and account for over 60% of the total debt outstanding as of 2002. All other variables are available for the universe of zip codes, but because house prices are the key right hand side variable in all of our analysis, we restrict ourselves to this sub-sample.

Much of the zip code level data is described in our earlier work (Mian, Rao, and Sufi (2013)), and we therefore relegate most of the discussion to the appendix. House price growth is measured using CoreLogic data. Income is available from the IRS Statistics of Income. We use a number of zip-code level variables from Equifax, including the fraction of subprime borrowers in a zip code. We use demographics from the 2000 Decennial Census, and the Saiz (2010) measure of housing supply elasticity, which is available at the CBSA (or metropolitan area) level. The only measure of household spending available to us at the zip code level is the quantity of new autos purchased by households living in a zip code, which comes from R.L. Polk and is based on vehicle registration information which lists the address of the person buying the vehicle, not the address of the dealer. In the appendix of Mian, Rao, and Sufi (2013), there is a table listing all of the data sources, the level of aggregation, and contacts for obtaining the data.

The methodology for constructing net worth, home values, and total dollars spent on new auto purchases in a zip code follows Mian, Rao, and Sufi (2013). We construct the dollar change in home values in a zip code by starting with the 2000 median home value estimate in the zip code according to the decennial Census multiplied by the number of owners in the zip code. We then grow this aggregate home value every year using house price growth in the zip code according to CoreLogic, and the change in homeownership rates for the country. For every year, we can then divide this zip code level home value by the number of households in the zip code, which gives us a home value per household in dollars for every year of our sample. The dollar value of auto purchases in a zip code is based on a proportionality assumption where we divide up aggregate expenditures on new autos into zip codes based on the proportion of autos purchased in the zip code. See the appendix for more details.

The only new data in this study are mortgage refinancing data come from CoreLogic. The underlying data are at the mortgage level and include information on outstanding mortgages, new mortgage issuances, and whether the new issuance is cash-out or no cash-out. The data used in this paper are aggregated at the zip code level. The CoreLogic data come from GSE's and large servicers. In terms of coverage it covers all GSE mortgages and about 70% of non-GSE mortgages. Unfortunately, we cannot measure in this data the actual number of dollars of equity taken out of homes because we do not see the outstanding principal on mortgages that are being refinanced. We only see the share of outstanding mortgage debt in a zip code refinanced in a cash-out transaction.

Given this shortcoming, for specifications on the marginal propensity to borrow, we utilize individual credit bureau data from Equifax used earlier in Mian and Sufi (2011). While the data provider had required earlier that we put individuals in groups of five for anonymity, we no longer face that constraint. Hence the Equifax data used in this paper is available at the individual level. The individual level sample is limited to homeowners, which are defined to be individuals who either had mortgage debt outstanding as of 1997, or individuals with zero mortgage debt outstanding but their credit report shows that they at some point had a mortgage. These data cover 60,856 homeowners as of 1997, and we track them through 2010. We repeat the details of this data in the appendix; they are also available in Mian and Sufi (2011). The individual level credit bureau data contain no information on spending.

Table 1 presents the summary statistics for the zip code level data, where all statistics are weighted by the number of households in the zip code as of 2000. The weighted statistics are relevant because all regressions include population weights. House price growth from 2002 to 2006 was on average 36%, with zip codes on average experiencing a $54.9 thousand increase in home equity value per household. A comparison of the 10th and 90th percentile reveals a large amount of variation across the country in house price growth during this period. In 2003 through 2006, an average of 10.5% of outstanding mortgages are refinanced in a cash-out transaction in a given year. The analogous number is 9.3% for no-cash-out refinancing. The change in the annual cash-out refinancing share from the 2003-2006 period relative to 2000-2002 period is 2.3%. The change is -1.7% for no-cash-out refinancing, which reflects the fact that interest rates rose from 2003 to 2006.

We use four different variables that are useful in separating zip codes by the propensity to borrow and spend according to cash on hand models: adjusted gross income, net worth, the fraction with a credit score below 660, or the fraction with less than a high school education. As we show below, all four variables are highly correlated.

Table 1 also presents information for the individual level homeowner data. The average change in debt for homeowners in the individual level sample was $52.1 thousand, and the average credit score in 1997 was 788.[4]

2. Theoretical Motivation and Estimation Strategy

2.1 Models of the effect of higher house prices on consumer behavior

Sinai and Souleles (2005) and Campbell and Cocco (2007) show that if homeowners are unconstrained then any increase in house prices makes future housing consumption more expensive. As a result, the propensity to spend out of housing gains is close to zero. However, housing gains also increase the homeowner's access to "cash on hand" if credit markets are willing to lend against higher collateral value. The cash-on-hand effect can be an important driver of household spending, especially for constrained households with low levels of wealth.

[4] The rise in debt in the individual sample from 2002 to 2006 is $52.1 thousand, whereas the rise in home value from 2002 to 2006 in the zip level sample is $54.9 thousand. This is obviously too large an increase in debt given the rise in home values. The explanation is that the individual level data only cover a subset of the overall sample: 1,988 zip codes covered by the Fiserv Case Shiller Weiss house price indices. These zip codes on average see a rise in home values of $144.8 thousand. In other words, FCSW covers a subset of zip codes with much higher home values. All zip level results are robust to isolating the sample to these FCSW zip codes.

The precautionary savings models of Deaton (1991) and Carroll and Kimball (1996) show that if income is uncertain and there is limited risk-sharing, then the consumption function is concave in wealth.[5] These models imply that cash-on-hand shocks translate into higher spending. More importantly, the concavity of the consumption function implies that the propensity to spend out of cash-on-hand shocks is higher for consumers with low levels of wealth. The heterogeneity in the marginal propensity to spend out of cash-on-hand shocks is further enhanced if one adds other reasons why households with low wealth might be credit constrained (see Carroll (2001) and cites therein).

A second strand of literature approaches this question from a different angle but arrives at the same conclusion. Harris and Laibson (2002) introduce time-inconsistent "beta-delta" preferences that lead to the following Euler equation:

$$u'(C_t) = R * E_t[\beta\delta C'(x_{t+1}) + \delta(1 - C'(x_{t+1}))]\, u'(C_{t+1})$$

where *x* represents cash on hand. The condition $\beta = 1$ implies that individuals are time consistent in their preferences, and the Euler equation reverts back to the standard Euler equation. Harris and Laibson (2002) computationally solve for the optimal consumption path in this framework, and show that $C(x_t)$ exhibits strong concavity: the marginal propensity to consume out of cash-on-hand shocks is larger for low cash-on-hand households (see their Figure 2). Given the beta-delta Euler equation, this result implies that households with low cash-on-hand act as if they have very high discount rate.[6]

2.2 Empirical strategy

[5] As Carroll and Kimball (1996) show, the third derivative of utility function needs to be positive, a condition satisfied by both CRRA and CARA utility functions.
[6] Kaplan and Violante (2014) introduce another rationale for high MPC out of cash-on-hand shocks based on a desire to hold otherwise illiquid but higher return assets. In Iacoviello (2005) rising housing values relieve a collateral constraint and therefore act as cash-on-hand shocks.

The most prominent empirical tests of the heterogeneity in MPC to cash-on-hand shocks come from the literature on fiscal stimulus rebates, and they find strong support for concavity. Households with lower cash on hand, measured as those with lower disposable income and financial wealth net of consumer debt, exhibit a stronger consumption response to rebate checks. Jappelli and Pistaferri (2014) find that the MPC out of rebate checks in Italy is 0.65 for the lowest cash-on-hand households, and 0.30 for the highest. Johnson, Parker, and Souleles (2006) find a more aggressive response by low income households to the 2001 stimulus checks in the quarter of receipt. [7]

The question we seek to answer is: Are housing wealth shocks also shocks to cash on hand? We answer this question by empirically estimating how changes in housing wealth affect household borrowing and spending decisions.[8] The object of interest is to estimate $\frac{\partial C(x_t)}{\partial x_t}$ separately for households of different types.

Our empirical specification exploits cross-sectional variation across zip codes in their exposure to housing wealth shocks from 2002 to 2006, and examines how these shocks affect both borrowing and consumption. The formal specification is:

$$\Delta y_{zc} = \alpha + \beta_1 * \Delta HomeValue_{zc} + \varepsilon_{zc} \qquad (1)$$

The right hand side variable is the dollar change in home value for a zip code z located in CBSA c from 2002 to 2006. The outcome variable Δy_{zc} is the dollar change in borrowing in Section 3, and the dollar change in auto purchases in Section 4, always measured contemporaneously from

[7] Hurst and Stafford (2002) ask a different question: can home equity be used to smooth consumption when a household faces income shocks? They find that the answer is yes.
[8] An increase in house prices in a zip code also represents an increase in the relative price of housing consumption service. However, we show below, the heterogeneity in response to house price changes seems much more consistent with the cash-on-hand framework rather than a model based on the relative price changes.

2002 to 2006. The specification is run in levels as opposed to logs to match the consumption function in the theoretical literature above.[9]

In equation (1), β_1 reflects the average marginal propensity to borrow or the average marginal propensity to consume (depending on the outcome variable) out of a dollar gain in home value. To test for concavity of the consumption function we need to sort zip codes by attributes, such as available cash on hand, that theory suggests lead to different MPCs. This leads to estimating:

$$\Delta y_{zc} = \alpha + \beta_1 * \Delta HomeValue_{zc} + \beta_2 * \Delta HomeValue_{zc} * CashonHand_{z,2002} + \beta_3 * CashonHand_{z,2002} + \varepsilon_{zc} \quad (2)$$

The key coefficient is the cross-derivative, β_2, which measures the effect on spending from home value shocks by the cash on hand of the zip code. A statistically significant negative value of β_2 would support models of cash on hand described above: high cash-on-hand households show a lower marginal propensity to borrow and consume out of housing wealth shocks.

There are two empirical challenges in estimating (1) and (2). The first challenge is ensuring that permanent income shocks are fixed when estimating the effect of rising home values on borrowing and spending. Indeed, as Carroll (2001) points out, computational solutions which produce a concave consumption function scale both cash on hand and consumption by permanent income. In our empirical setting, the biggest worry is that shocks to house prices are *positively correlated* with unobservable permanent income shocks. Such a positive correlation would lead us to spuriously associate the consumption response to the home value shock, when it instead is due to a change in permanent income of the household. The second challenge is

[9] In this study, we do not specify the exact source of the *aggregate* shock that led to higher demand for housing. We have argued in our previous research that the shock was due to an increase in the supply of credit (Mian and Sufi (2009, 2011)), and this argument has been supported by other research (Landvoigt, Piazzesi, and Schneider (2013), Favilukis, Ludvigson, and Van Nieuwerburgh (2013)). In this study, we exploit cross-sectional variation in house price growth across the country coming from the aggregate demand shock for housing.

measuring $CashonHand_{z,2002}$. What zip code level characteristic should we use to measure the theoretical notion of cash on hand in the models described above?

For the first challenge, we follow our earlier work and instrument home value changes with the housing supply inelasticity of the city in which the zip code or individual is located. The leads to the following two stage least squares estimation:

$$\Delta y_{zc} = \alpha^{IV} + \beta_1^{IV} * \Delta \widehat{HomeValue}_{zc} + \beta_2^{IV} * \Delta \widehat{HomeValue_{zc} * CashonHand_{z,2002}} + \beta_3^{IV} * CashonHand_{z,2002} + \varepsilon_{zc} \tag{3}$$

$$\Delta HomeValue_{zc} = \omega + \eta_1 * Inelasticity_c + \eta_2 * Inelasticity_c * CashonHand_{z,2002} + \eta_3 * CashonHand_{z,2002} + \epsilon_{zc} \tag{3a}$$

$$\Delta HomeValue_{zc} * CashonHand_{z,2002} = \psi + \lambda_1 * Inelasticity_c + \lambda_2 * Inelasticity_c * CashonHand_{z,2002} + \lambda_3 * CashonHand_{z,2002} + \zeta_{zc} \tag{3b}$$

Equations 3a and 3b represent the first stage specifications where the instruments are housing supply inelasticity of a city and inelasticity interacted with cash on hand in the zip code as of 2002. The endogenous variables are home value change from 2002 to 2006 and the interaction of home value change from 2002 to 2006 with cash on hand in the zip code as of 2002. Estimation of equation 3 produces the coefficient β_2^{IV}, the instrumental variables estimate of concavity of the consumption function. We also report non-parametric instrumental variables specifications where we split zip codes into four income groups. In these specifications, the first stage estimation instruments are inelasticity interacted with each of the four income groups.

For the second challenge, we follow the extant research and utilize measures of income in the zip code in 2002 as a measure of cash on hand. However, as we show in the appendix, the results are robust to using net worth per household in a zip code or credit scores in the zip code. These variables are all highly correlated, as we show below.[10]

2.3 First stage and exclusion restriction

The left panel of Figure 1 shows the evolution of house prices in the United States from 1999 to 2010. House prices rose from 1999 to 2002 at a steady rate, and then accelerated substantially from 2002 to 2006. For the right panel, we sort zip codes into population-quartiles based on the housing supply elasticity of the city in which they are located. The right panel plots the highest and lowest quartile. House prices rose significantly more in inelastic housing supply CBSAs relative to elastic housing supply CBSAs, which shows the power of the first stage.[11]

Columns 1 and 2 of Table 2 present the zip-code level first stage estimation where the left hand side variable is the log difference and level difference in house prices from 2002 to 2006, respectively. For now, we shut down the heterogeneity channel. The estimate in column 1 implies that a one standard deviation increase in inelasticity is associated with a 12% increase in house prices from 2002 to 2006, which is more than a half standard deviation. Column 2 implies that a one standard deviation increase in inelasticity is associated with a $15 thousand increase in home value, which is a 30% standard deviation.

The housing supply inelasticity instrument has substantial power in predicting house price growth. What about the exclusion restriction? In column 3, we repeat a result from Mian and Sufi (2011). We regress the wage growth shock from 2002 to 2006 in a zip code on the

[10] The main specifications reported in the text do not include control variables. In the appendix, we use a number of controls and show that none of the main results are affected.
[11] In the appendix, we show the scatter-plot of house price growth from 2002 to 2006 against housing supply inelasticity. We also show the scatter-plot of residential investment against housing supply inelasticity.

housing supply inelasticity instrument. As mentioned above, the wage growth shock in a zip code from 2002 to 2006 represents wage growth per household in a zip code from 2002 to 2006 subtracting the wage growth in the zip code from 1998 to 2002. We see an estimate of almost exactly zero, with a small confidence interval. Wage growth does not accelerate differentially in inelastic CBSAs from 2002 to 2006.[12]

Our primary focus in the tests below is the heterogeneity across the income distribution of the effect of house prices on borrowing and spending. As a result, a stronger test of the exclusion restriction is to see whether *low income* zip codes in inelastic housing supply CBSAs saw differential *positive* wage growth shocks from 2002 to 2006. Column 4 shows evidence of the opposite: *higher* income zip codes saw acceleration of wage growth from 2002 to 2006. This is consistent with the well documented rise in inequality during the decade. In column 5, we interact the income of the zip code with the inelasticity of the CBSA. The point estimate on the interaction term suggests that high income zip codes in inelastic CBSAs had somewhat lower relative wage growth versus the high income zip codes in elastic CBSAs, but the estimates have a large standard error and we cannot reject that the interaction term is zero.

Under the assumption that observable acceleration of wage growth from 2002 to 2006 is an accurate measure of permanent income shocks, we see no evidence that low income zip codes in inelastic housing supply cities saw positive permanent income shocks. In fact, the evidence suggests the opposite. Lower income zip codes saw negative permanent income shocks relative to high income zip codes during our sample period.

Table 3 reports the correlation across zip codes in various measures of cash on hand. All three measures we consider--income per household, net worth per household, credit scores, and

[12] Table II of Mian, Rao, and Sufi (2013) shows that housing supply elasticity is uncorrelated with the 2006 employment share in construction, construction employment growth from 2002 to 2006, and population growth from 2002 to 2006.

education levels--are highly correlated. While education levels are less directly related to cash on hand, they may help capture differences in *beta-delta* type behavior, which is why we include them in Table 3. In the appendix, we show that our results are robust to the use of the alternative measures of cash on hand. The last row of Table 3 shows that the wage growth shock from 2002 to 2006 is smaller in low cash on hand shock zip codes no matter how we measure them.

3. The Marginal Propensity to Borrow out of Housing Gains

3.1. Zip-code level results

Table 4 presents zip-code level regressions of the share of mortgages refinanced with cash taken out on house price growth. The left hand side variable in these regressions is the average annual share of mortgages refinanced in a cash-out transaction from 2003 to 2006 less the average annual share of mortgages refinanced in a cash-out transaction from 2001 to 2002. Unfortunately, in zip code-level data, we do not know the exact number of dollars borrowed against rising home values because we cannot see the principal balance of the mortgage refinanced. As a result, we cannot estimate an exact marginal propensity to borrow out of a dollar increase in home values. We can do so in individual data which we report in the next subsection. For now, we provide qualitative evidence by showing the relation between cash-out refinancing share in a zip code and house price growth.

Column 1 presents the OLS regression of the change in cash-out mortgage refinancing share on house price growth. There is a large positive effect. The point estimate implies that a one standard deviation increase in house price growth (20%) leads to a 3 percentage point increase in the annual share of mortgages refinanced with cash out, which is 2/3 a standard deviation of the left hand side variable. Columns 2 and 3 show heterogeneity in this effect across the 2002 zip-code level income distribution. The result is much weaker in higher income zip codes. The magnitudes are easiest to gauge in the non-parametric specification in column 3 where we split zip codes into four groups. For the highest income zip codes, those with average household AGI of $100 thousand or higher, the effect is (0.092/0.152 =) 60% weaker.

As mentioned above, the OLS specification suffers from the concern that house price movements are positively correlated with permanent income movements. The housing supply inelasticity instrument can help in this regard. The left panel of Figure 2 reports the reduced form version of the instrumental variables specification, focusing on zip codes with less than $50 thousand in per-household income. Cash-out refinancing was almost identical in inelastic and elastic housing supply CBSAs in 2001. Zip codes in inelastic CBSAs refinanced slightly more in 2002, but then see a large relative jump from 2003 to 2006. After 2006, the difference collapses. By the end of the sample period, cash-out refinancing is identical for the two groups.

The right panel of Figure 2 shows that no-cash-out refinancing, or refinancing driven only by lower interest rates, is identical for zip codes in inelastic or elastic CBSAs throughout the sample period. The right panel suggests that these two groups have similar loadings on refinancing propensity driven by credit supply frictions or other factors. The difference in refinancing seems uniquely related to cash-out refinancing in response to higher house prices.

Columns 4 through 6 of Table 4 show the IV estimates. The results are similar to the OLS results. The one difference worth noting is in column 6. In the IV estimation, the cash-out refinancing share response to house price growth for the highest income zip codes has a point estimate of almost zero (0.177-0.164 = 0.013). When house prices rise, households in high income zip codes do not undertake a cash-out refinancing.

In Table 5, we use the dollar change in home value in a zip code from 2002 to 2006 as the right hand side variable instead of house price growth. Qualitatively, the results are almost identical. The estimate in column 1 implies that a $50 thousand increase home values (one standard deviation) leads to a 3 percentage point increase in the annual share of mortgages refinanced in a cash-out transaction, which is about 2/3 a standard deviation. We again find that this effect is substantially weaker in high income zip codes. In the non-parametric specifications in column 3, the effect of home value changes on the cash-out refinancing share is again close to zero for zip codes with per household income of $100 thousand or higher.

Unlike Table 4, Table 5 shows a difference between the OLS and IV specifications. The average effect in the IV estimates is twice as large as in the OLS (column 1 versus column 4). In both the OLS and IV results, higher income zip codes are much less responsive to higher home values. But the effect of a dollar increase in home value on cash-out refinancing share is much larger for low income zip codes according to the IV estimate. We return to an explanation of this result in the next subsection.

3.2 Individual level results

In Table 6, we utilize individual-level credit bureau data on homeowners from Equifax. These are individuals that already owned their home as of 1997. Table 6 reports the estimated coefficients from equations (1), (2), and (3) developed in Section 2.3 above. The main difference is that level of observation is an individual rather than a zip code. The OLS estimate in column 1 implies a $0.09 per dollar marginal propensity to borrow out of increases in housing wealth. The IV estimate in column 4 is twice as large, showing a marginal propensity to borrow out of housing wealth of $0.19 per dollar.[13]

[13] Our estimate in Mian and Sufi (2011) was $0.25 per dollar. The sample and estimation are almost identical, and so the difference is entirely due to the fact that we put individuals into groups of five in the previous study.

Why are the IV estimates so much larger in the dollar on dollar specifications? We believe it is because the OLS estimates are biased downward. Recall the evidence from Table 3 above that low income zip codes experience worse wage growth shocks than high income zip codes from 2002 to 2006. Local variation in house price growth, or variation across neighborhoods in a city with the same housing supply elasticity, is partly driven by these differential wage growth shocks. Therefore, there is likely to be a spurious positive correlation between the 2002 income level in a zip code and the house price shock at the local within-city level. Given that the true treatment effect on borrowing is much lower for homeowners with higher income, the spurious positive correlation biases the OLS coefficient downward. The IV specification in contrast does not rely on such local variation, and instead uses across-city variation in housing supply elasticity. Table 3 shows that housing supply elasticity is uncorrelated with wage growth shocks, and therefore provides the cleaner source of variation.

The estimates in columns 1 and 4 are average marginal propensities to borrow. We are interested in heterogeneity across the cash-on-hand distribution. Columns 2 and 5 sort individuals based on their zip-code level income, as in Tables 4 and 5. In both columns 2 and 5, we see a negative coefficient estimate on the interaction term, but it is statistically weak especially in the IV specification.

However, the weak statistical power of the interaction term is to be expected given that we do not observe individual level income directly. To see this point, let V_{iz} be the variance of the interaction term $(X_z * W_{iz})$, where X_z is change in home value in zip code *z* and W_{iz} is the income of individual *i* in *z*. Since we do not observe W_{iz} directly, we are forced to use a proxy for an individual's income using the zip code level average income $\overline{W}_z$. Let $\overline{V}_z$ be the variance of the interaction term $(X_z * \overline{W}_z)$ that is used in the actual regression. For simplicity, assume X_z and

W_{iz} are independent. It follows that $\bar{V}_z < V_{iz}$. The standard error of the coefficient is inversely proportional to the variance of the interaction term, and therefore the standard error will blow up whenever we are forced to take averages for the interaction variable.

We therefore use the credit score of an individual in 1997 as an alternative interaction variable since this variable is measured at the individual level. We don't have income at the individual level, but we do have credit scores. We already know from Table 3 that credit scores and income are highly correlated at the zip code level, and so credit scores can also be interpreted as a measure of cash on hand. For example, Mian and Sufi (2009) show that credit scores are a powerful predictor of whether credit applications are denied, and Gross and Souleles (2002) use credit card utilization rates--which are very highly correlated with credit scores--as a measure of liquidity constraints.

The estimated coefficients in columns 3 and 6 on the interaction term of credit scores and house price growth are negative and statistically significant in both the OLS and IV specifications. To get a sense of magnitudes, Figure 3 provides estimates of the marginal propensity to borrow across the credit score distribution. It uses an estimation similar to the one reported in column 6 of Table 6, except we include four credit score bins non-parametrically in the estimation, instead of the linear credit score variable. The marginal propensity to borrow is more than \$0.25 per dollar for individuals with a credit score below 700. It is just over \$0.20 per dollar for individuals between 700 and 799. There is a large drop off for individuals between 800 and 899, and the effect is zero for individuals with a credit score above 900. In terms of the distribution, 22% of the homeowners in our sample have a credit score below 700. For the next three categories, the corresponding numbers are 26%, 38%, and 14%, respectively.

The results in Table 6 and Figure 3 show a striking degree of heterogeneity across the population in the marginal propensity to borrow against increases in housing wealth.[14] Lower credit score, lower income households treat a rise in home equity as a cash-on-hand shock. They aggressively liquefy home equity through cash-out mortgage refinancing. Rising home equity may affect consumption because it leads to cash on hand for low cash-on-hand individuals. In the next section, we examine whether these same individuals increase spending on new autos.

4. The Marginal Propensity to Spend out of Housing Gains

4.1 Effect on auto purchases

Table 7 reports estimated coefficients from specifications (1), (2), and (3) outlined in Section 2.3. The outcome variable is the change in the dollar amount spent on new auto purchases per household in 2006 less the amount spent in 2002. The estimate in column 1 implies a marginal propensity to spend on autos out of housing wealth of 1.6 cents per dollar. The estimate is about half a cent smaller than the result from 2006 to 2009 reported in Mian, Rao, and Sufi (2013). Notice that this is not a *cumulative* estimate of the impact of home value changes from 2002 to 2006 on auto purchases from 2002 to 2006. It is the estimate for purchases in 2006 alone. We report the cumulative effect later in this sub-section.

The estimate in column 1 is an average effect. Columns 2 and 3 show strong heterogeneity in the effect across the 2002 zip-code level income distribution. The coefficient estimate on the interaction term in column 2 is negative and statistically significant at the 1 percent level. The estimates in column 3 imply an MPC on autos out of housing wealth of 2.5 cents per dollar for households living in zip codes where the average income is less than $35

[14] This is a new result relative to Mian and Sufi (2011). In the previous study, we showed that the *elasticity* of borrowing with respect to house prices was stronger for low credit score homeowners, but we never showed the differences in the marginal propensity to borrow out of increases in home equity wealth.

thousand. For households living in zip codes where average income is $100 thousand or more, the MPC is (0.025 - 0.018 =) 0.7 cents per dollar.

The IV estimates are similar to the OLS estimates.[15] In column 6, we find complete offset of the MPC effect among zip codes living in the highest income zip codes. So households living in zip codes with average income of $35 thousand or less have an MPC on autos of 2.6 cents per dollar. Households living in zip codes with average income of more than $100 thousand have an MPC of almost exactly zero. Figure 4 plots the heterogeneity in MPCs across the zip code income distribution.[16]

As mentioned above, the estimates in Table 7 are for the effects of home value changes from 2002 to 2006 on auto purchases in 2006 only. In Figure 5, we estimate the cumulative effect using the following methodology. We fix the right hand side variable to be the home value change from 2002 to 2006, instrumented using housing supply inelasticity. We then estimate the effect of the home value change from 2002 to 2006 on the change in auto purchases for years 2000 through 2006. For each year, this gives us an estimate of the incremental autos purchased in that year due to the housing effect. The estimate for 2006 is already reported in column 4 of Table 7. We report 2000 and 2001 to get a sense of any pre-trend in the data. Once we have the coefficient estimates of the incremental impact for each year, we can add them to get the cumulative effect from 2002 to 2006.

[15] We had discussed in Section 3 how a spurious within-city correlation between home value changes and unobserved income growth leads to a downward bias for the OLS estimate of marginal propensity to borrow relative to the IV estimate. The same logic however does not apply when the dependent variable is changed to auto purchases. The reason is that with auto purchases there is an additional bias in the OLS specification and that bias goes in the opposite direction. In particular, while zip codes with unobserved higher income growth shocks do not borrow against their homes, they likely increase their spending on durables such as automobiles in response to wage growth shocks. It appears that the net effect of these biases makes OLS and IV similar. But the IV remains our preferred estimate.

[16] Given that autos are often purchased with loans, especially among lower income individuals, one question is whether households are furthering leveraging their home equity withdrawals with auto debt. In Section 4, we estimate the effect of home value changes on total debt which includes auto loans, and so it already includes this further leveraging.

Figure 5 shows the estimates. There is no evidence of a pre-trend from 2000 to 2002. The effect of home value changes on auto purchases is small in 2003, but then jumps significantly in 2004 through 2006. As mentioned above, the bar for 2006 gives the estimate of 0.017 which is already in Table 7. If we add the bars from 2003 to 2006, we get a cumulative effect of 4.4 cents per dollar of home equity change. So from the end of 2002 to the end of 2006, homeowners spent 4.4 cents of every dollar rise in home values on new autos.

4.2 Using alternative structural break instrumental variable

The empirical strategy above relies on housing supply elasticity as an instrument. A recent paper by Charles, Hurst, and Notowidigdo (2014) proposes a new instrument for house price growth from 2002 to 2006 that is meant to capture the "bubble" component of house price appreciation.[17] The instrument is based on the assumption that underlying fundamental factors do not change abruptly, and evidence of a sharp break in the house price growth process in a zip code or city is evidence of a bubble unrelated to fundamentals. The bigger the structural break, the larger is the value of this "bubble" instrument.

We follow the exact procedure outlined in Charles, et al (2014) and construct estimates of structural break in the house price series at the quarterly zip code level.[18] In particular, we use data between 2000Q1 and 2005Q4 and search for zip code-specific structural break by searching for the location of break that maximizes the R-sq of the following regression:

$$\log(P_{zt}) = \alpha_z + \beta_z * t + \gamma_z * (t - t_z^*) * 1\{t > t_z^*\} + \epsilon_{zt} \qquad (4)$$

where P_{zt} is house price index for zip code *z* in quarter *t* and γ_z represents the structural break that maximizes R-sq of (4) by searching over t_z^* between 2002Q1 and 2004Q4. We use the

[17] The authors motivate their instrument based on work by Ferreira and Gyourko (2011), Chinco and Mayer (2014), and Glaeser, Gyourko, and Saiz (2008) that argues that pure speculative activity can lead to short-run bubbles exemplified by structural breaks.
[18] The results are very similar if we instead define the instrument at the CBSA level.

estimated γ_z, which measures the size of the structural break—the "bubble"—as an instrument for home value changes between 2002 and 2006. We re-run all of our empirical tests using this alternative instrument and Appendix Table 3 summarizes the results for our key regressions.

The results show a strong first stage. But unlike the housing supply elasticity instrument, the structural break instrument *is* strongly correlated with the wage growth shock. Charles, et al (2014) argue that this wage growth shock was not a permanent income change by showing that employment suffered when house prices collapsed. In other words, the exclusion restriction assumption of Charles et al (2014) is that there was no structural break in fundamentals even though the wage bill increased in these areas. Under this assumption, we can proceed with the IV estimates. Columns 4 through 6 of Appendix Table 3 show that there is strong heterogeneity across the income distribution in the marginal propensity to borrow and spend. The effect of house price growth on spending is twice as strong in low income zip codes using the Charles, et al (2014) instrument, which may reflect the fact that households in "bubble" cities perceived the temporary earnings boost as permanent.

4.3 Autos versus all other spending

The estimated marginal propensity to spend only includes spending on new automobile purchases given that this is the only spending measure we have at the zip code level from 2002 to 2006. However, we can impute the marginal propensity to spend on all goods using results from Mian, Rao and Sufi (2013).

Mian, Rao, and Sufi (2013) use *county-level* total spending measure from 2006 to 2009 to estimate the marginal propensity to spend out of housing wealth loss between 2006 and 2009.[19] They estimate a marginal propensity to consume per dollar of housing wealth loss of $0.054 for total spending and of $0.023 for new auto purchases. Auto spending constitutes 42.6% of the

[19] These county level data are not available prior to 2006.

overall MPC. If we assume that auto spending constitutes the same percentage of overall spending at the margin from 2002 to 2006, then our cumulative MPC for auto spending of 4.4 cents per dollar suggests that the total marginal propensity to consume out of the rise in house values from 2002 to 2006 is (0.044*(0.054/0.023)=) $0.10 per dollar. [20]

Two other studies also provide some guidelines. Parker, Souleles, Johnson, and McClelland (2013) analyzed the 2008 stimulus payments which were on the order of $300 to $600 for individuals and $600 to $1200 for families, and they find a significant effect on new auto purchases. They do not provide a precise estimate of the marginal propensity to spend on autos, but they find that households spent 12-30 percent of the stimulus payments without considering durables, and 50-90 percent with durables. Aaronson, Agarwal, and French (2009) find that modest increases in the minimum wage lead to vehicle purchases. They find that a $300 per quarter increase in family income following a $1 minimum wage hike leads to about an $800 increase in spending on vehicles.

How does the $0.10 per dollar estimate for the overall MPC out of housing wealth compare with the marginal propensity to borrow out of housing wealth? The marginal propensity to borrow was $0.19 per dollar. However, this estimate cannot be directly compared with the overall MPC because the $0.19 per dollar estimate is conditional on the borrower being a homeowner, while the $0.10 per dollar spending response is estimated at the zip code level and therefore includes both homeowners and renters. The effect of higher house prices on auto spending for renters is likely to be zero, and may even be negative given that they are likely to be increasing their consumption of housing services going forward.

[20] Mian, Rao, and Sufi (2013) do not provide estimates of the heterogeneity in MPCs for all goods across the distribution because county-level data is too aggregated to have power for detecting heterogeneity.

We can impute the MPC for homeowners under the assumption that the MPC for renters is zero. In our sample, 63% of households are homeowners. The implied cumulative MPC for homeowners is 0.10/0.63=0.16, or $0.16 per dollar. Our estimate is reasonably close to the independent estimate of the marginal propensity to borrow for homeowners, $0.19 per dollar. If we assume that the MPC of spending for renters is slightly negative, then these two numbers will line up even closer.[21]

There are two important implications of this finding. First, almost all of the money borrowed out of home equity is used for spending, where spending is broadly defined to include residential investment such as home improvement.[22] Second, almost all of the spending out of home equity is driven by borrowing. In other words, the housing wealth effect is better described as a housing borrowing effect--households increase their spending when home values rise because it allows for borrowing. This is consistent with the cash-on-hand framework we discussed in Section 2.

5. Macroeconomic Implications

How large is the *aggregate* impact of house price appreciation between 2002 and 2006 on consumer spending given the estimated MPC out of housing gains? The answer to this important question depends on the overall gain in home values between 2002 and 2006 that can be considered "exogenous" and general equilibrium feedback effects. We address each of these issues below.

[21] One may worry that the sample of zip code in the individual level data is a strict subsample of the sample of zip codes used in Table 7 to estimate the MPC. Hence the MPC estimate that corresponds to the individual level data zip codes may be different. However, our estimate of MPC is almost identical when we restrict Table 7 to zip codes that are also included in the individual level data.

[22] The Mian, Rao, and Sufi (2013) estimate includes retail spending on all goods, which includes local purchases of goods that would be used in home improvement.

5.1 Quantifying the total effect on spending driven by house price appreciation

The results above show that the MPC out of housing wealth varies by a zip code's income level. It is therefore necessary to take into account the heterogeneity of the treatment effect across income levels when computing the total effect of housing wealth on spending.

Let *I* index one of the four income categories of a zip code used in Table 7. We can compute the aggregate impact of the change in home value on household spending by computing $\Phi = \sum_z (\beta_I^{IV} * \Delta HV_{z,I} * Pop_{z,I})$, where β_I^{IV} is the instrumental variables estimate of the MPC for zip codes in income category *I*, $\Delta HV_{z,I}$ is the change in home value (in dollars) for zip code *z* belonging to income category *I*, and $Pop_{z,I}$ is the number of households in zip code *z* belonging to income category *I*. Given that our sample of zip codes represent 55.4% of total auto sales in 2002, we divide the computed sum by 0.554 to get an estimate that is representative of total population.

A key question in the computation of Φ is how to calculate the change in home value, $\Delta HV_{z,I}$. The MPC estimate β_I^{IV} is obtained using cross-sectional variation in house price changes across CBSAs that is orthogonal to changes in income or fundamentals. By using the actual gain in home value from 2002 to 2006 in the aggregate calculation, we are implicitly assuming that all of the house price gains in the U.S. economy from 2002 to 2006 are independent of fundamental improvements in income or productivity.

We believe there is a good case to be made for this assumption. The 2002 to 2006 period represents a time when the price to rent ratio deviated wildly form its long run mean, and by 2010 it had returned to its 2002 level.[23] Most of the deviation in prices between 2002 and 2006 is likely to represent a deviation from fundamentals. However, our aggregate calculation can easily

[23] See, for example, http://www.calculatedriskblog.com/2013/05/real-house-prices-price-to-rent-ratio.html.

be adjusted under alternative scenarios where only a fraction of overall growth in home values is considered to be orthogonal to productivity or permanent income shocks.

Our methodology yields an estimate for the total boost in spending due to house price gains between 2002 to 2006 equal to 0.08% of GDP in 2003, 0.8% of GDP in 2004, and 1.3% of GDP in both 2005 and 2006. Thus the cumulative boost in spending between 2002 and 2006 is equal to 3.5% of GDP.

5.2 Possible general equilibrium effects

The calculation above suggests a shift in total spending, or aggregate demand, of 3.5% of GDP between 2002 and 2006. One can think of this as a shift in the aggregate demand curve, all else equal. We now explore whether this shift in aggregate demand translates into an *equilibrium* increase in output and demand. In order to do so, we need to incorporate our estimate of the shift in spending into a general equilibrium model. The purpose of this model is to provide some basic guidelines on what general equilibrium feedback effects might offset the increase in spending coming from the rise in home values.

One of our key findings is that there are two distinct types of consumers in the United States, the less wealthy "hand-to-mouth" consumers who have a very high propensity to borrow and spend, and wealthy consumers who do not respond much to the same home equity shocks. The simplest way to model this finding is to imagine an economy with two types of agents, "borrowers" (B) and "lenders" (L).

Consider an overlapping generations model in the spirit of Eggertsson and Mahotra (2014). A new generation of borrowers and lenders is born each period and lives for two periods as young (y) and old (o). Let Y_t denote total output at time t, a fraction $(1-g)$ of which goes to the borrowers when they get *old* and g goes to the lenders when they are *young*. The asymmetric income timing implies that the borrowers who are young want to borrow from lenders who are young in every period.

We assume that borrowers when young are always credit constrained and can only borrow up to their debt limit D. Borrowers' consumption is then given by $C_t^{B,y} = \frac{D}{(1+r_t)}$ when they are young and $C_{t+1}^{B,o} = (1-g)Y_{t+1} - D$ when they are old, where r_t is the real interest rate.

Assuming log preferences and perfect foresight, lenders that are young have an inter-temporal tradeoff given by their Euler equation, $\frac{1}{C_t^{L,y}} = \beta(\frac{1+r_t}{C_{t+1}^{L,o}})$. Let B_t be the total savings of lenders when young that is loaned to the young borrowers. Then $C_t^{L,y} = gY_t - B_t$ and $C_{t+1}^{L,o} = B_t(1+r_t)$. Solving these equations simultaneously gives us the aggregate loan supply function, $B_t = \frac{\beta}{(1+\beta)} gY_t$.

Loan market equilibrium implies that loan supply equals loan demand, which we know is $\frac{D}{(1+r_t)}$. This gives us the steady state relation between the real interest rate and other variables:

$$1 + r_t = \frac{(1+\beta)D}{\beta g Y_t} \qquad (5)$$

An increase in the share of income going to the lenders or an increase in lenders' patience lowers the real interest rate. An increase in borrowing capacity of borrowers increases the real interest rate.

We can introduce nominal prices in the economy by using the Fisher equation, $1 + r_t = \frac{(1+i_t)}{(1+\pi_t)}$, and assuming that monetary policy follows the Taylor principle given by $(1 + i_t) = Max\{1, (1+i^*)\left(\frac{1+\pi_t}{1+\pi^*}\right)^\tau\}$, for some target $i^*, \pi^*, and\ \tau > 1$.

Combining (5) with the Fisher equation and the Taylor principle gives us the steady state aggregate demand (AD) curve. Assuming that the nominal interest rate does not hit the zero lower bound, the downward sloping AD curve is given by, $\frac{(1+i^*)(1+\pi_t)^{\tau-1}}{(1+\pi^*)^{\tau}} = \frac{(1+\beta)D}{\beta g Y_t}$. Taking logs and approximating $log(1+\pi_t)$ with π_t, we can write down the AD curve as:[24]

$$\pi_t = \gamma^* + \frac{d}{(\tau-1)} - \frac{y_t}{(\tau-1)} \tag{6}$$

We close the model by specifying a traditional aggregate supply (AS) curve of the form:

$$y_t = \bar{y} + \kappa\pi_t \tag{7}$$

Where $\kappa(.)$ is an increasing function of wage rigidity, and $\kappa = 0$ corresponds to the benchmark case of no nominal rigidity and hence a vertical AS curve. Increasing wage rigidity implies a flatter AS curve with $\kappa > 0$ – the classic Phillips curve. Equilibrium output is determined by the intersection of the AD and AS curves as shown in Figure 6, which plots deviation from steady state output on the horizontal axis and deviation from steady state inflation on the vertical axis.

An increase in home values increases the debt capacity of constrained consumers (the borrowers) from D to $D' > D$, and increases spending by Δy, which we have estimated to be 3.5% of GDP between 2002 and 2006. This shifts out the AD curve from AD1 to AD2 as shown in Figure 6, *all else equal*. The equilibrium impact of the shift in AD curve on output depends on the shape and possible shift in the AS curve. We consider these possibilities below.

First, consider the situation where the aggregate supply curve is vertical, labelled AS1 in Figure 6. In this case, the shift in AD curve results in *no* aggregate change in spending because the entire increase in aggregate demand is absorbed by higher inflation – the equilibrium moves from point A to point B. In other words, the increase in spending due to the shift in aggregate demand curve caused by higher debt capacity is completely offset by consumers forced to reduce their demand due to higher inflation.

[24] Where $\gamma^* = \frac{1}{(\tau-1)} log\left(\frac{(1+\pi^*)^{\tau}}{(1+i^*)}\frac{(1+\beta)}{\beta g}\right)$ and lower case letters correspond to logs of respective uppercase letters.

In general the upward pressure on inflation is an important metric through which we can gauge the strength of the general equilibrium feedback effect through the AS curve. Using the calibration of $\tau = 2$ for the Taylor principle as in Eggertsson and Mehrotra (2014), the implied *increase* in inflation necessary to keep total output the same is 3.5 percentage points.[25] This is much larger than the actual change in inflation during the 2002 to 2006 period.

The left panel of Figure 7 plots the PCE core inflation index (excluding food and energy) over time. We exclude food and energy because commodities tend to be more volatile over shorter horizons, and their prices are driven more by international trade forces than domestic demand. Core inflation hovered between 1.4% and 2.4% during the 2002 to 2006 period with inflation being on the higher end of this range towards the end of the period. So while there is slight tendency for inflation to increase, there is no evidence for an increase in inflation sufficient to fully negate the 3.5% of GDP shift in aggregate demand.

Moreover, overall price inflation may not be the proper inflation metric to look at given that the marginal propensity to spend out of housing gains is heavily skewed toward durable goods, and in particular toward the purchase of new automobiles. The right panel of Figure 7 plots the CPI for new vehicles and shows that there is no evidence of any price pressure for new automobiles between 2002 and 2006. Inflation for new vehicles is between -2% and 1% during this period. Observed inflation from 2002 to 2006 suggests an absence of a general equilibrium price effect that would negate the spending shift in aggregate demand due to housing gains.

The muted, if any, response of inflation to the shift in spending suggests that either the AS curve is flat, or that our sample period was also accompanied by a shift in the AS curve. A flatter AS curve (labelled AS2) can go some distance in explaining why inflation does not increase by as much as 3.5 percentage points (point B versus B’). However, the typical

[25] Coibion and Gorodnichenko (2011) estimate a long-run Taylor rule coefficient of 2.5 for inflation.

calibration of new Keynesian Phillips curve implies that the AS curve is not flat enough to generate the low inflation response observed in the data.

As an alternative, we consider the possibility that the AS curve also shifted out during the period – shown as AS3 in Figure 6. One interpretation of the shift in the AS curve is the increased productivity in the outside world – particularly in China.[26] The ability of the external trade account to absorb most of the increase in domestic spending may help explain why the shift in aggregate spending did not result in an upward pressure on inflation. A shift in AS curve moves the equilibrium from point A to point C, implying that most of the increase in spending due to the AD shift is sustained in equilibrium.[27]

Another possibility relates to the appropriate counter-factual for our general equilibrium thought experiment. A number of commentators have recently resurrected Alvin Hansen's "secular stagnation" idea that aggregate demand in developed world might be facing strong headwinds, leading to a continuous decline in the real interest rate and a tendency for the AD curve to shift inwards (Summers (2013)). As *The Economist* puts it, in such a scenario "advanced economies will keep inflating bubbles in a doomed attempt to resurrect growth".[28] Under this view, in the absence of a housing bubble between 2002 and 2006, the AD curve would have shifted inwards. Relative to this counter-factual, our estimated aggregate demand shift represents a net increase in output even without any material shift in the AS curve.

6. What Happens Afterward?

[26] We thank Kenneth Rogoff for pushing us in this direction.
[27] Of course the increase in output in this example is shared with China.
[28] See *The Economist*, http://www.economist.com/news/finance-and-economics/21591181-old-explanation-economic-drift-gains-new-following-stagnant-thinking

Models which predict a concave consumption function have different explanations for the behavior of low cash-on-hand households. In models of liquidity constraints and precautionary savings, a high MPC out of cash-on-hand shocks by low cash-on-hand households reflects a high marginal utility of consumption today because of credit market imperfections that disallow households from borrowing against future income. More cash today relieves the pressure on the household to self-insure through precautionary savings, and so higher spending today is unambiguously a good thing. In behavioral models such as the *beta-delta* framework, high consumption today is suboptimal from the perspective of the prior self who would instead want to commit the excessively-impatient present self from engaging in such behavior.

Distinguishing these models is beyond the scope of our analysis. But some descriptive evidence may help us make progress. In particular, what happens to these zip codes that aggressively borrowed and consumed out of housing wealth shocks after the housing boom? Table 8 presents estimated coefficients that help answer this question. In particular, the specifications reported regress measures of wage growth and auto spending growth on the housing supply inelasticity instrument and the instrument interacted with each of the income categories. This is the "reduced form" of these outcomes on the instruments.

Column 1 shows that wages fall from 2006 to 2011 the most in low income zip codes located in inelastic CBSAs.[29] These are exactly the zip codes that borrow and spend the most aggressively during the housing boom. Column 2 looks at the wage growth shock from 2006 to 2011, which is wage growth from 2006 to 2011 minus wage growth from 2002 to 2006. Once again, we see that low income zip codes located in inelastic CBSAs have the worst wage growth shocks from 2006 to 2011.

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[29] We use wage growth from 2006 to 2011 instead of 2006 to 2009 because IRS data are not available for 2009 or 2010.

Columns 3 and 4 examine auto sales. Whether we use auto sales growth or the decline in dollars spent on autos, we see the same result: spending declines the most from 2006 to 2009 in exactly the zip codes that borrowed and spent the most. This is perhaps less surprising given the results in Mian, Rao, and Sufi (2013) showing that low income zip codes cut back spending the most for a given decline in house prices from 2006 to 2009. Further, given that a new auto is a durable good, we should not be surprised that spending declines most in areas that have the newest stock of autos.

Table 8 shows that households that borrow and spend the most during the housing boom experience the largest decline in income and spending during the housing bust. The income result is difficult to reconcile with a standard liquidity constraints model in which households borrow when constraints are loosened because they believe future income will be higher. It could be that the negative shock to wages was an unlikely outcome in the *ex ante* beliefs of households living in these zip codes. But it is telling that aggressive borrowing predicts lower wage growth and lower consumption in the future.

7. Conclusion

The rise in U.S. house prices from 2002 to 2006 had a large effect on household spending. We provide evidence on why house prices affected spending, and we quantify how large the effect was during the U.S. housing boom. House prices affected spending because low cash-on-hand households treated the rise in home values as a cash-on-hand shock. In particular, households living in low income zip codes aggressively liquefied home equity in response to rising house prices. On average, homeowners borrowed $0.19 per $1 of home equity gains from 2002 to 2006. There is strong heterogeneity in this effect, with the lowest cash-on-hand households withdrawing $0.25 per dollar, and the highest cash-on-hand households being unresponsive.

Households living in low income zip codes spent this extra cash. We find that on average, a $1 rise in home values from 2002 to 2006 led to a $0.04 increase in new auto purchases. Once again, there is strong heterogeneity. Households living in high income zip codes were completely unresponsive. We are able to attribute almost all of the effect of house prices on spending to the borrowing channel. House prices matter for spending because they facilitate borrowing by low income households. We quantify the aggregate effect of house prices on household spending using a number of assumptions. Our estimate is that rising home values added 0.08% to GDP in 2003, 0.8% in 2004, and 1.3% in both 2005 and 2006.

Why exactly do low income households aggressively borrow and spend? This question is extremely important, but also extremely challenging to answer. We show that households in the zip codes that most aggressively borrowed and spent out of rising home values *ex post* had *lower* income growth and spending. While the *ex post* realization is one of only many potential *ex ante* outcomes, it does suggest some skepticism of the view that households in these zip codes expected high future income. We hope that future research can address why low income households borrow and spend so aggressively out of house price shocks.

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Figure 1
House Prices

This figure plots house prices for the United States from 1999 to 2010. The left panel shows house prices for the country as a whole. The right panel shows house prices for the top and bottom population-weighted quartile Core-Based Statistical Areas (CBSAs, or cities) based on the housing supply elasticity measure of Saiz (2010).

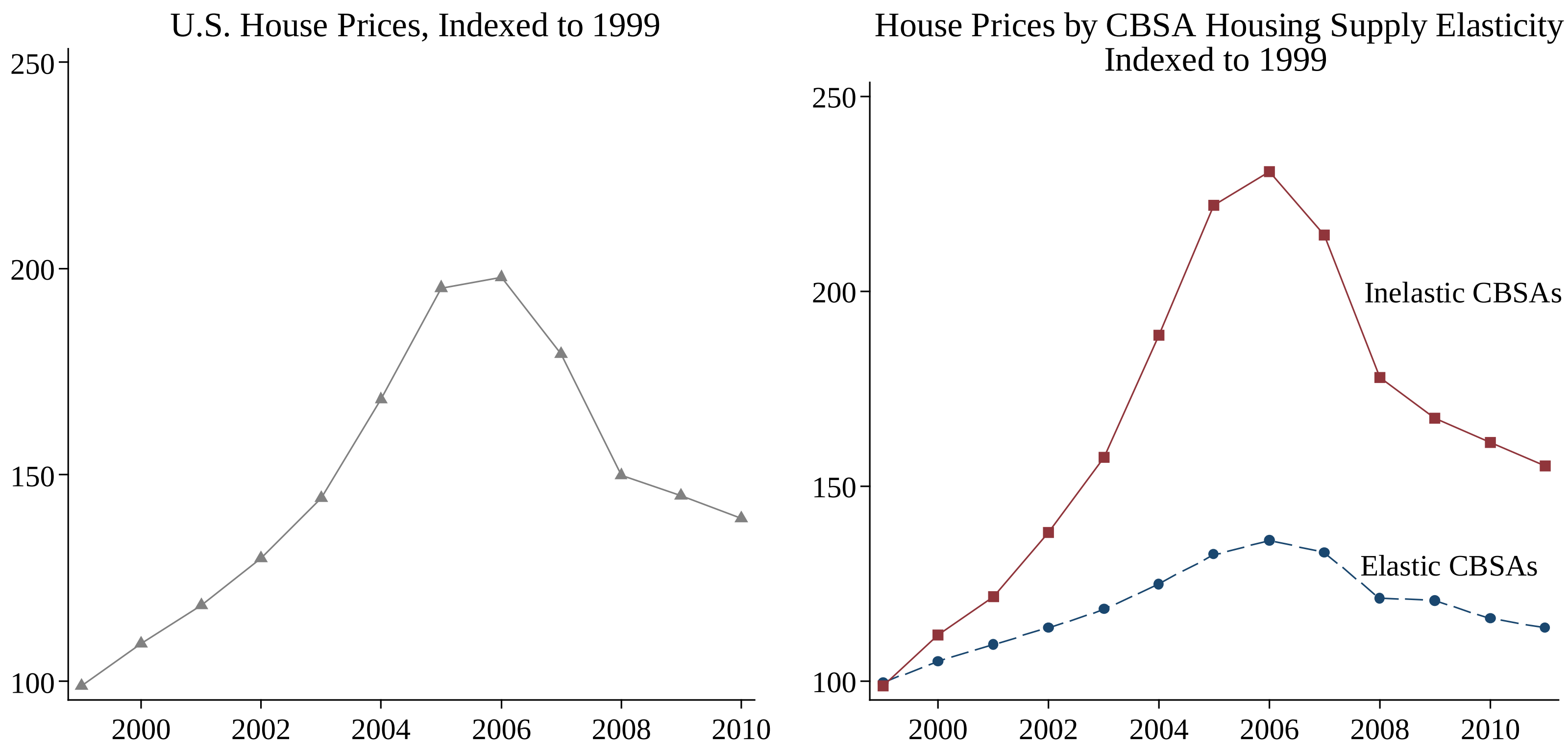

Figure 2
Mortgage Refinancing for Low Income Zip Codes

This figure plots the share of outstanding mortgages refinanced in a given year for the top and bottom population-weighted quartile Core-Based Statistical Areas (CBSAs, or cities) based on the housing supply elasticity measure of Saiz (2011). The sample is limited to zip codes where average per household income is $50 thousand or less. The left panel plots refinancing where cash was taken out, and the right panel plots refinancing where no additional cash was taken out.

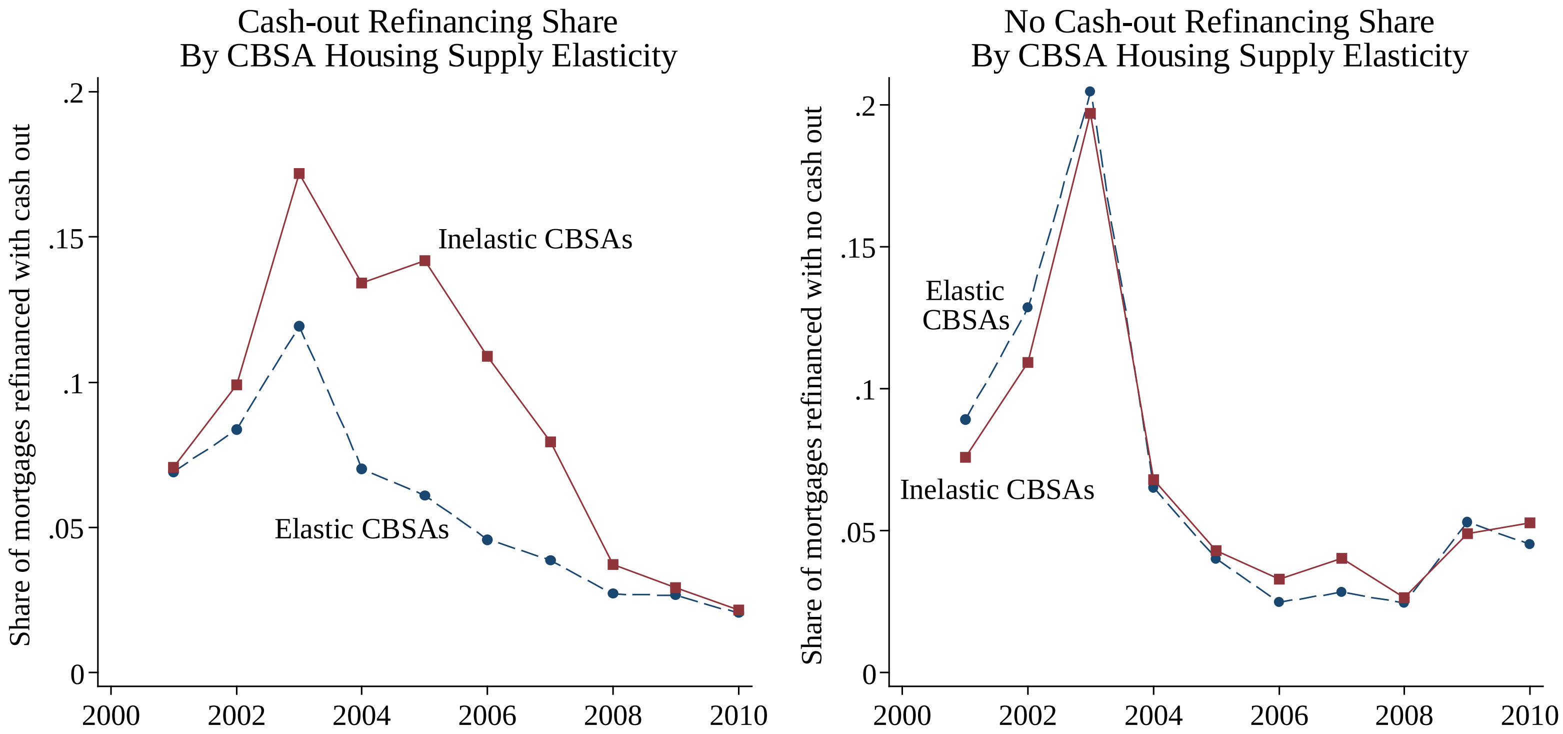

Figure 3
Heterogeneity in Marginal Propensity to Borrow out of Housing Wealth

This figure plots the marginal propensity to borrow out of a $1 increase in home equity value, by credit score of the household. The credit score we use is the Vantage Score. The underlying estimation uses individual level data and examines the effect of the rise in home values from 2002 to 2006, instrumented with housing supply inelasticity, on total borrowing.

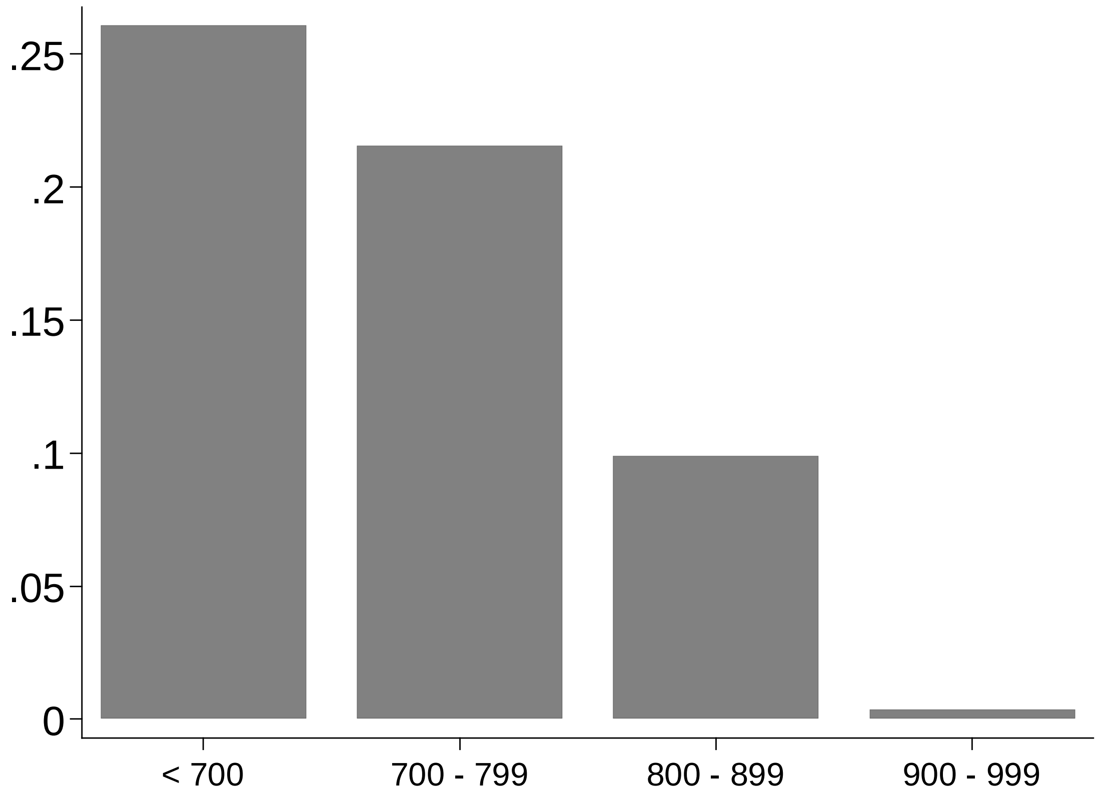

Figure 4

Heterogeneity in Marginal Propensity to Spend on Autos out of Housing Wealth

This figure plots the marginal propensity to spend on new autos out of a $1 increase in home equity value, by per-household adjusted gross income (AGI) of a zip code. The estimation use zip code level data and examines the effect of the rise in home values from 2002 to 2006, instrumented with housing supply inelasticity, on increase in purchases of new autos in 2006 relative to 2002.

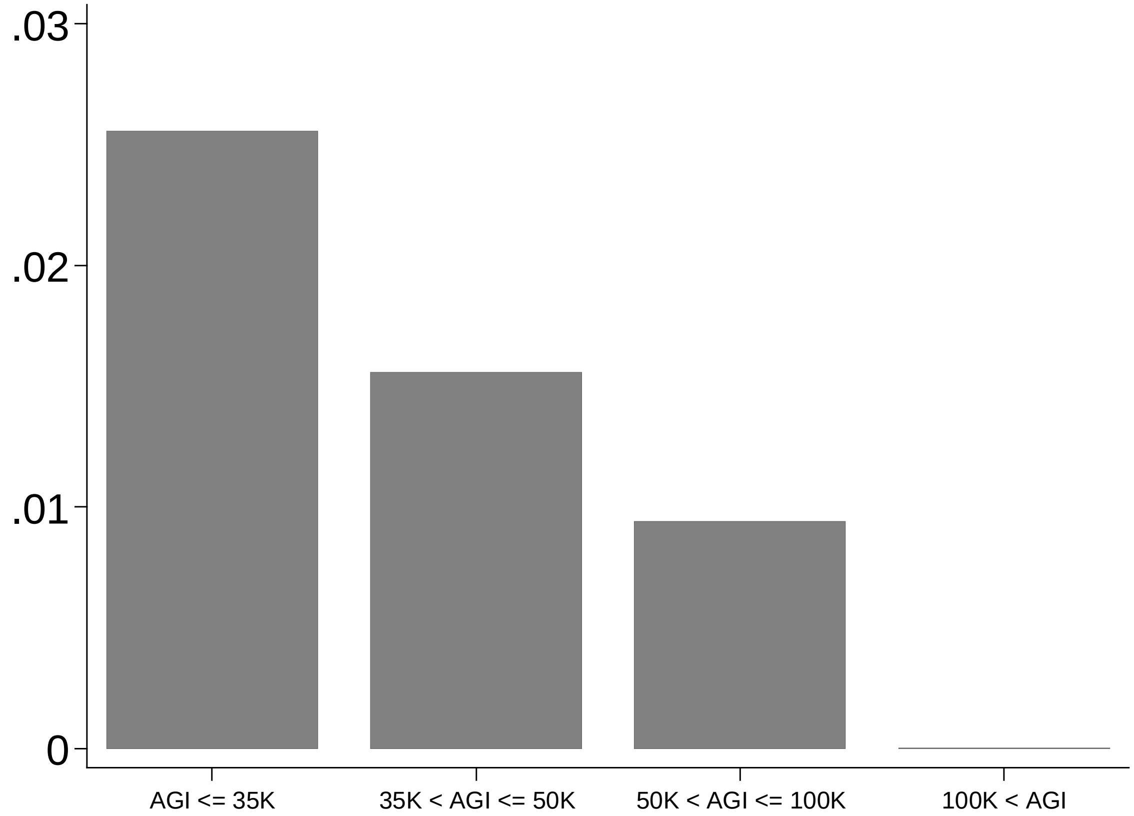

Figure 5
Marginal Propensity to Spend out of Housing Wealth by Year

This figure plots the marginal propensity to spend on new autos out of a $1 increase in home equity value by year. We fix the right hand side variable to the be the dollar change in home equity value from 2002 to 2006, and we plot the estimated effect on the dollar change in new autos from 2002 to each year on the *x*-axis. Adding up these bars from 2002 to 2006 gives the cumulative effect of the house price increase on auto spending.

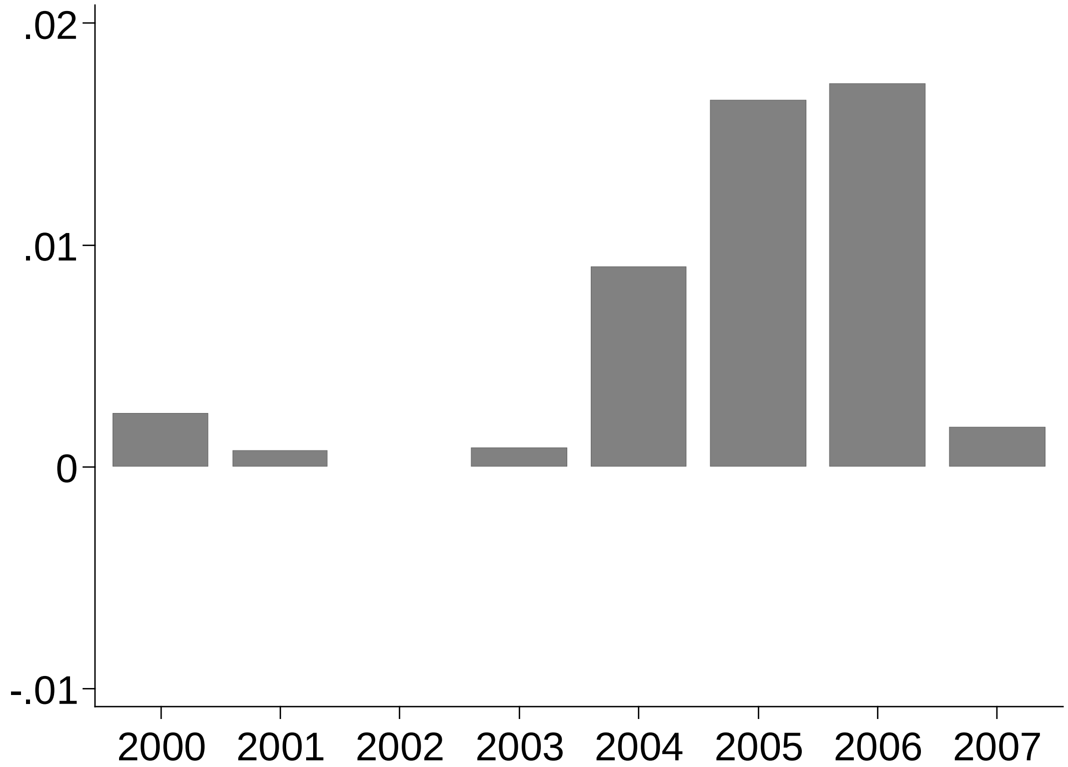

Figure 6
The Effect of Aggregate Demand Shift in General Equilibrium

This figure plots the aggregate demand (AD) and aggregate supply (AS) relation derived in the text. The y-axis plots deviation of inflation from initial steady state, and the x-axis plots deviation of log output from initial steady state. Under the assumptions highlighted in text, the increase in house prices between 2002 and 2006 shifts the aggregate demand curve by 3.5% of GDP from AD1 to AD2. As a result of this shift in aggregate demand, the new equilibrium can shift from point A to either point B, B’ or C depending on the shape of the AS curve, and any shift in the AS curve itself. As the graph shows, the equilibrium impact on inflation is one of the key factors that can help us select between the three possibilities. See text for a complete discussion.

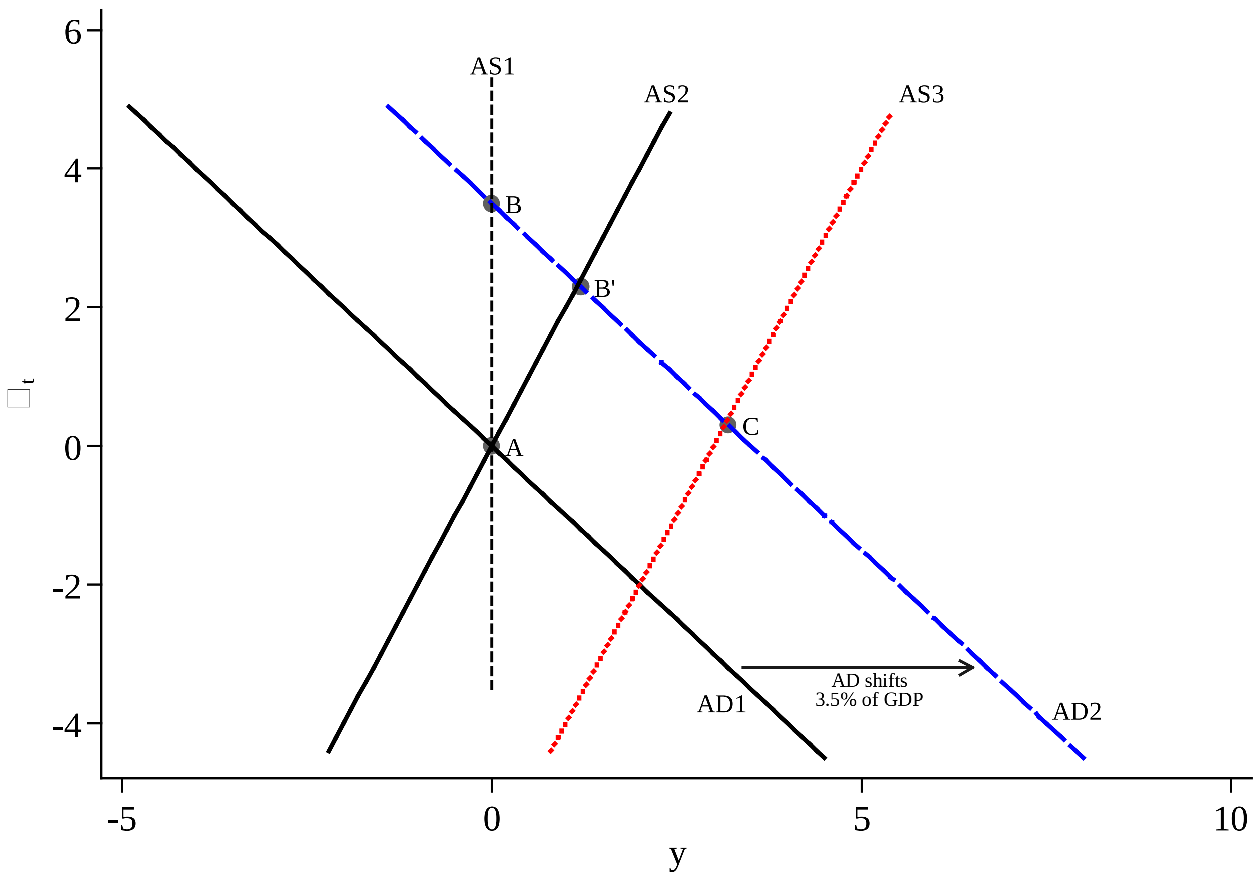

Figure 7
Inflation during Housing Boom

This figure plots the year-on-year growth in PCE core inflation and New Vehicles inflation over time. For comparison, 2% and 2.5% inflation benchmarks are represented by dashed and dotted lines respectively. The shaded area represents the 2004-2006 period when the estimated spending against rising house prices is the strongest.

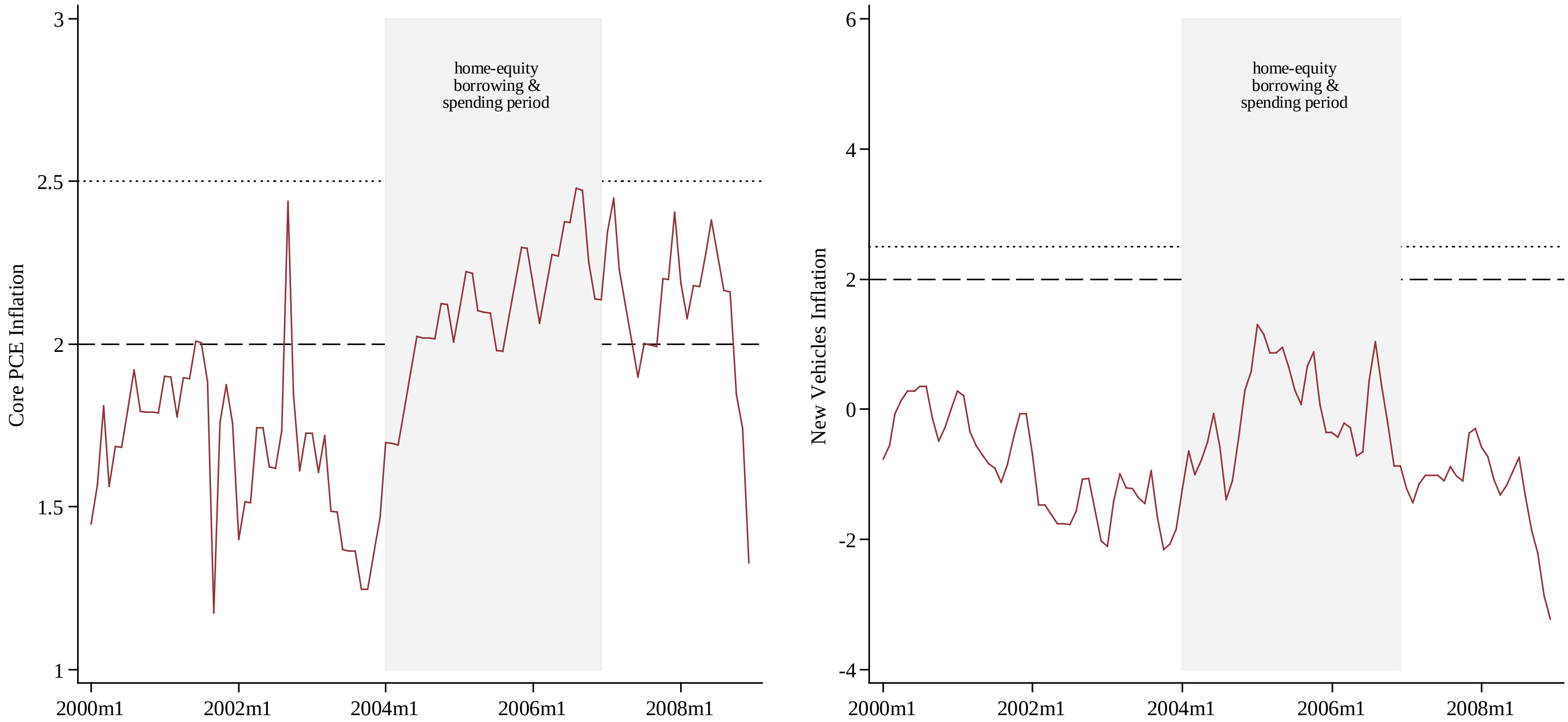

Table 1
Summary Statistics

This table presents zip code level and individual level summary statistics for our sample, which runs from 2000 to 2010. Housing supply inelasticity is based on the Saiz (2010) measure. Our measure is (5-elasticity)/5 in order to make it vary between 0 and 1.The cash-out and no-cash-out refinancing shares are the annual flow of cash-out and no-cash-out refinancing scaled by outstanding mortgages at the beginning of the year. The wage growth shock from 2002 to 2006 is wage growth in the zip code from 2002 to 2006 minus wage growth in the zip code from 1998 to 2002. All summary statistics are weighted by the number of households in the zip code as of 2000.

	N	Mean	SD	10th	90th
Zip level data					
House price growth, 2002 to 2006	5163	0.360	0.217	0.089	0.658
Change in home value ($000), 2002 to 2006	5163	54.9	54.0	6.6	121.8
Annual cash-out refinancing share, 2003 through 2006	5163	0.105	0.047	0.049	0.170
Annual no-cash-out refinancing share, 2003 through 2006	5163	0.093	0.032	0.055	0.133
Change in annual cash-out refinancing share	5163	0.023	0.043	-0.024	0.080
Change in annual no-cash-out refinancing share	5163	-0.017	0.020	-0.043	0.004
Change in new auto purchases per household, ($000), 2002 to 2006	5163	0.862	3.129	-1.056	2.655
Housing supply inelasticity	5163	0.672	0.179	0.404	0.865
Adjusted gross income per household ($000), 2002	5163	49.9	25.1	29.5	74.5
Net worth per household ($000) 2002	5163	322.4	303.6	108.5	597.2
Fraction with credit score below 660, 2002	5163	0.344	0.134	0.180	0.531
Less than high school education fraction, 2000	5163	0.173	0.110	0.056	0.321
Wage growth shock, 2002 to 2006	5162	-0.019	0.083	-0.093	0.058
Median home value ($000), 2002	5163	176.7	109.5	83.2	305.0
Number of households, thousands	5163	14.4	6.7	6.9	22.7
Individual level homeowner data					
Change in debt ($000), 2002-2006	60858	52.1	203.4	-69.2	229.2
Credit score, 1997	60858	788.4	103.4	641.0	910.0

Table 2
First Stage and Exclusion Restriction

Columns 1 and 2 of this table present the first stage estimation of house price growth and home value change on the Saiz (2010) housing supply elasticity instrument. Our inelasticity measure is (5-elasticity)/5 in order to make it vary between 0 and 1. The wage growth shock from 2002 to 2006 is wage growth in the zip code from 2002 to 2006 minus wage growth in the zip code from 1998 to 2002. All regressions are weighted by the number of households in the zip code as of 2000. Standard errors are clustered at the CBSA level.

	(1) House price growth, 2002-2006	(2) Home value change ($000), 2002-2006	(3) Wage growth shock 2002-2006	(4) Wage growth shock, 2002-2006	(5) Wage growth shock, 2002-2006
Housing supply inelasticity	0.673**	86.720**	0.001		0.009
	(0.089)	(13.824)	(0.017)		(0.037)
Median home value, 2002 ($000)		0.256**			
		(0.048)			
AGI per household ($000), 2002				0.647**	0.925
				(0.101)	(0.717)
Inelasticity*AGI per household ($000)					-0.370
					(0.908)
Constant	-0.092	-48.574**	-0.020	-0.052**	-0.059*
	(0.048)	(10.049)	(0.012)	(0.005)	(0.027)
Observations	5,163	5,163	5,162	5,162	5,162
R-squared	0.310	0.484	0.000	0.038	0.039

** p<0.01, * p<0.05

Table 3
Correlations between Cash-on-Hand Measures

The table presents correlations between different measures of cash on hand across zip codes. The wage growth shock from 2002 to 2006 is wage growth in the zip code from 2002 to 2006 minus wage growth in the zip code from 1998 to 2002. All correlations are statistically distinct from zero at the 1% level of confidence.

	(1) AGI per household ($000), 2002	(2) Net worth per household ($000), 2002	(3) Fraction with credit score below 660, 2002	(4) Fraction with less than high school education, 2000
Net worth per household ($000), 2002	0.849			
Fraction with credit score below 660, 2002	-0.610	-0.525		
Fraction with less than high school education, 2000	-0.556	-0.409	0.679	
Wage shock, 2002 to 2006	0.079	0.114	-0.094	-0.077

Table 4
The Effect of House Price Growth on Cash-out Refinancing Share

This table presents regressions of the change in the cash-out refinancing share in a zip code on house price growth. The change in cash-out refinancing share is the average annual share of mortgages refinanced in a cash-out transaction from 2003 to 2006 minus the average annual share in 2001 and 2002. Regressions are weighted by the number of households in the zip code as of 2000. Standard errors are clustered at the CBSA level.

	(1)	(2)	(3)	(4)	(5)	(6)
	Change in cash-out refinancing share, 2002 to 2006					
	OLS	OLS	OLS	IV	IV	IV
House price growth, 2002-2006	0.142**	0.178**	0.152**	0.128**	0.192**	0.177**
	(0.010)	(0.017)	(0.012)	(0.012)	(0.026)	(0.014)
(HP growth, 02-06)*(AGI, 2002)		-0.854**			-1.281*	
		(0.268)			(0.529)	
AGI per household ($ millions), 2002		-0.041			0.116	
		(0.105)			(0.151)	
(HP growth, 02-06)*($35K < AGI < $50K)			-0.009			-0.035**
			(0.010)			(0.013)
(HP growth, 02-06)*($50K < AGI < $100K)			-0.038*			-0.102**
			(0.015)			(0.021)
(HP growth, 02-06)*(AGI > $100K)			-0.092**			-0.164**
			(0.019)			(0.043)
$35K < AGI < $50K			-0.006*			0.004
			(0.003)			(0.004)
$50K < AGI < $100K			-0.009			0.014*
			(0.005)			(0.006)
AGI > $100K			-0.004			0.023
			(0.008)			(0.013)
Constant	-0.028**	-0.024**	-0.020**	-0.023**	-0.029**	-0.030**
	(0.003)	(0.006)	(0.004)	(0.004)	(0.007)	(0.004)
Observations	5,163	5,163	5,163	5,163	5,163	5,163
R-squared	0.526	0.579	0.587	0.520	0.577	0.571

** p<0.01, * p<0.05

Table 5
The Effect of Home Value Changes on Cash-out Refinancing Share

This table presents regressions of the change in the cash-out refinancing share in a zip code on home value changes. The change in cash-out refinancing share is the average annual share of mortgages refinanced in a cash-out transaction from 2003 to 2006 minus the average annual share in 2001 and 2002. Regressions are weighted by the number of households in the zip code as of 2000. Standard errors are clustered at the CBSA level.

	(1)	(2)	(3)	(4)	(5)	(6)
	Change in cash-out refinancing share, 2002 to 2006					
	OLS	OLS	OLS	IV	IV	IV
Home value change ($000), 2002-2006	0.0006**	0.0007**	0.0011**	0.0014**	0.0014**	0.0017**
	(0.0001)	(0.0001)	(0.0001)	(0.0002)	(0.0001)	(0.0002)
(HV change, 02-06)*(AGI, 2002)		-0.0024**			-0.0064**	
		(0.0006)			(0.0023)	
AGI per household ($ millions), 2002		-0.2944**			0.2048	
		(0.1026)			(0.2038)	
(HV change, 02-06)*($35K < AGI < $50K)			-0.0003**			-0.0006**
			(0.0001)			(0.0001)
(HV change, 02-06)*($50K < AGI < $100K)			-0.0007**			-0.0011**
			(0.0001)			(0.0001)
(HV change, 02-06)*(AGI > $100K)			-0.0010**			-0.0012**
			(0.0001)			(0.0002)
$35K < AGI < $50K			-0.0104**			0.0027
			(0.0030)			(0.0028)
$50K < AGI < $100K			-0.0068			0.0120**
			(0.0045)			(0.0042)
AGI > $100K			0.0192*			0.0207
			(0.0077)			(0.0194)
Median home value, 2002	-0.0002**	-0.0001**	-0.0001**	-0.0005**	-0.0003**	-0.0002**
	(0.0000)	(0.0000)	(0.0000)	(0.0001)	(0.0000)	(0.0000)
Constant	0.0306**	0.0292**	0.0168**	0.0317**	0.0057	0.0065
	(0.0051)	(0.0067)	(0.0039)	(0.0088)	(0.0080)	(0.0036)
Observations	5,163	5,163	5,163	5,163	5,163	5,163
R-squared	0.301	0.387	0.517		0.163	0.405

** p<0.01, * p<0.05

Table 6
The Marginal Propensity to Borrow out of Housing Wealth
Evidence from Individual-Level Data on Homeowners

This table presents individual level regressions of the change in total debt on home value change from 2002 to 2006. Debt is measured at the individual level. The home value change is measured at the zip code level. AGI per household in 2002 is measured at the zip code level, whereas the credit score is measured at the individual level. Standard errors are clustered at the CBSA level.

	(1)	(2)	(3)	(4)	(5)	(6)
	Change in total debt ($000), 2002 to 2006					
	OLS	OLS	OLS	IV	IV	IV
Home value change ($000), 2002-2006	0.088**	0.143**	0.575**	0.188**	0.206**	0.797**
	(0.023)	(0.042)	(0.156)	(0.049)	(0.053)	(0.168)
(HV change, 02-06)*(AGI, 2002)		-0.971*			-1.014	
		(0.417)			(0.664)	
AGI per household ($ millions), 2002		2.827			96.155	
		(94.207)			(113.301)	
(HV change, 02-06)*(Credit score, 1997)			-0.063**			-0.082**
			(0.018)			(0.018)
Credit score (divided by 100), 1997			-3.513			-0.490
			(2.079)			(1.906)
Median home value, 2002	0.120**	0.159**	0.161**	0.063*	0.112*	0.122**
	(0.012)	(0.025)	(0.013)	(0.026)	(0.053)	(0.024)
Constant	13.338**	4.976	33.765	12.209**	1.947	8.700
	(3.168)	(4.710)	(18.617)	(3.981)	(6.337)	(17.012)
Observations	60,856	60,856	60,856	60,856	60,856	60,856
R-squared	0.012	0.012	0.016	0.010	0.012	0.016

** p<0.01, * p<0.05

Table 7
The Effect of Home Value Changes on New Auto Purchases

This table presents regressions of the change in dollars spent on new auto purchases from 2002 to 2006 on the change in home values during the same period. Regressions are weighted by the number of households in the zip code as of 2000. Standard errors are clustered at the CBSA level.

	(1)	(2)	(3)	(4)	(5)	(6)
	Change in New Auto Purchases ($000), 2002 - 2006					
	OLS	OLS	OLS	IV	IV	IV
Home value change ($000), 2002-2006	0.016**	0.022**	0.025**	0.017**	0.027**	0.026**
	(0.003)	(0.003)	(0.005)	(0.005)	(0.007)	(0.008)
(HV change, 02-06)*(AGI, 2002)		-0.082**			-0.177**	
		(0.016)			(0.051)	
AGI per household ($ millions), 2002		13.219**			23.890**	
		(3.366)			(6.693)	
(HV change, 02-06)*($35K < AGI < $50K)			-0.005			-0.010
			(0.004)			(0.006)
(HV change, 02-06)*($50K < AGI < $100K)			-0.010*			-0.016*
			(0.004)			(0.007)
(HV change, 02-06)*(AGI > $100K)			-0.017**			-0.026**
			(0.005)			(0.009)
$35K < AGI < $50K			0.093			0.264
			(0.191)			(0.239)
$50K < AGI < $100K			0.442			0.685*
			(0.254)			(0.305)
AGI > $100K			1.893**			2.417**
			(0.413)			(0.746)
Median home value, 2002	-0.004**	-0.005**	-0.005**	-0.005**	-0.005*	-0.003
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
Constant	0.762**	0.204	0.507*	0.764**	-0.345	0.291
	(0.143)	(0.180)	(0.195)	(0.135)	(0.278)	(0.205)
Observations	5,163	5,163	5,163	5,163	5,163	5,163
R-squared	0.043	0.049	0.051	0.043	0.041	0.046

** p<0.01, * p<0.05

Table 8
Income and Spending After the Housing Boom

This table presents regressions of wage growth and the growth in spending on new autos to the housing supply inelasticity of the CBSA in which the zip code is located. The wage growth shock from 2006 to 2011 is wage growth in the zip code from 2006 to 2011 minus wage growth in the zip code from 2002 to 2006. Regressions are weighted by the number of households in the zip code as of 2000. Standard errors are clustered at the CBSA level.

	(1) Wage growth, 2006-2011	(2) Wage growth shock, 2006-2011	(3) Growth in new auto purchases, 2006 to 2009	(4) Change in New Auto Purchases ($000), 2006 - 2009
Housing supply inelasticity	-0.073*	-0.133**	-0.233**	-2.787**
	(0.030)	(0.031)	(0.060)	(0.704)
(Inelasticity)*($35K < AGI < $50K)	0.054*	0.057**	0.040	0.608
	(0.023)	(0.022)	(0.042)	(0.570)
(Inelasticity)*($50K < AGI < $100K)	0.122**	0.093**	0.123*	1.712*
	(0.031)	(0.035)	(0.050)	(0.753)
(Inelasticity)*(AGI > $100K)	0.143*	0.136	0.146	5.047*
	(0.066)	(0.130)	(0.105)	(2.366)
$35K < AGI < $50K	-0.002	-0.013	0.036	-0.536
	(0.015)	(0.013)	(0.025)	(0.369)
$50K < AGI < $100K	-0.017	-0.022	0.038	-1.088*
	(0.019)	(0.023)	(0.030)	(0.492)
AGI > $100K	-0.029	-0.129	0.063	-3.080
	(0.049)	(0.100)	(0.073)	(1.808)
Constant	0.050**	0.006	-0.301**	-0.341
	(0.018)	(0.018)	(0.033)	(0.372)
Observations	5,162	5,162	5,163	5,163
R-squared	0.104	0.049	0.149	0.019

** p<0.01, * p<0.05