A New Perspective on the Relationship Between House Prices and Income

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Draft: 16 June 2008

We show that a strong linear relationship exists between income and house price quantiles in Sydney (Australia), Houston, and the state of Texas. This suggests that the house price distribution is closely approximated by the income distribution after a location-scale transformation. The slope of the line changes over time in response to changes in the mortgage market. We argue that this finding is consistent with a simple variant on the permanent income hypothesis. We then explore some of the implications with regard to the evolution of house prices, price-to-income ratios, the efficiency of the housing market, the construction and interpretation of hedonic price indexes for housing, and for public policy. (*JEL.* C43, E01, E31, G12, R31)

Keywords: Real Estate; Permanent income; Mortgage market; Housing bubble; Hedonic index

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1 Introduction

House prices and price-to-income ratios have risen significantly over the last decade in many Western countries. This has led to debate over whether this rise has been caused by a bubble, particularly in the financial press [see for example Woodall (2003)]. The academic literature has been more circumspect in its analysis of this topic. These trends it is argued could also be attributed to factors such as falling interest rates, demographic shifts and supply constraints.

In this paper we explore the relationship between income, house prices and these other factors. Our starting point is a remarkable empirical regularity that we have uncovered between income and house prices in Sydney (Australia), Houston, and the state of Texas. We find that an almost perfect linear relationship exists between income and house price quantiles for every year we consider. The slope of the line varies over time. The strength of this linear relationship and the fact that it holds in two very different cities and even at state level in the US leads us to conjecture that it may hold more generally. If true, it implies that house price distributions are closely approximated by their corresponding income distributions after a location-scale transformation.

We show how a linear relationship between income and house prices arises naturally from a variant on the permanent income hypothesis. In this setting, house prices can rise either due to movements along the income-house price line or due to a shift of the line. Upward shifts in the line are caused, for example, by a fall in the mortgage interest rate, deregulation of the mortgage market, and tougher zoning laws.

We conclude by exploring some of the implications of our findings for the evolution of house prices, price-to-income ratios, efficiency of the housing market, hedonic price indexes for housing, and for public policy.

2 Linking House Prices to Income

Our objective is to explore the relationship between the distribution of house prices and the distribution of income. We do this using data sets for Sydney, Houston and

Texas. Our data set for Sydney, Australia consists of house price sales for 198 postcodes over the period 1996-2006, and gross household income data for the years 1996, 2001 and 2006. Our housing data set was obtained from Australian Property Monitors (APM). We trim the top and bottom 0.5 percent of the price data due to the greater concentration of data entry errors in the tails of the distribution. For the years 1996, 2001 and 2006, respectively, we are left with a total of 58163, 86915 and 66808 house price sales observations. Our income data were obtained from the Australian Bureau of Statistics (ABS) Census. The house price distribution for each of these years is graphed in Figure 1 using the Sheather-Jones plug-in kernel density estimation method. Some summary statistics are provided in Table 1. The mean, median, standard deviation, skewness, and kurtosis all rise over time.[1]

Insert Figure 1 Here

Insert Table 1 Here

We are not able to match up income and house prices across households. Instead, we compare quantiles. That is, we match up the qth quantile of the house price distribution with the qth quantile of the income distribution. A plot of house prices (Y_{tq}) against gross income (X_{tq}) for t equals 1996, 2001 and 2006 are shown in Figure 2.

Insert Figure 2 Here

The quantile plots for all three years reveal a strong linear relationship between house prices and income.[2] This finding is confirmed by our regression results. We run cross-section regressions for 1996, 2001 and 2006 across matched decile pairs of Y_{tq} and X_{tq}:

$$Y_{tq} = \alpha_t + \beta_t X_{tq} + \varepsilon_{tq}. \tag{1}$$

Our results are shown in Table 2. The R^2 coefficients of 0.99, 0.99 and 0.95 are particularly striking, thus confirming a clear linear relationship.

Insert Table 2 Here

[1]The Jarque-Bera test rejects the null hypothesis that the original data or log transformed data are drawn from Gaussian distributions. If follows that it is not appropriate to assume that the house price data has a log-normal distribution.

[2]There is a slight suggestion of nonlinearity in 2006.

The US housing market differs in a number of important ways from that in Australia. In particular, in Australia only investors can tax deduct mortgage interest payments [see Bourassa (1996)] while in the US only owner occupiers can [see Green and Wachter (2005)]. Also, mortgages are typically fixed rate in the US and variable rate in Australia. In spite of these differences, it will be shown that a strong linear relationship also exists between income and house price quantiles in Houston, and indeed for the state of Texas itself.

Our data set for Houston consists of house price sales and gross household income data over the period 1999-2006. Our housing data set was obtained from the Real Estate Center at Texas A & M University and the income data from the American Community Survey (ACS). We only have access to a frequency distribution of house prices, not the raw data itself. Nevertheless, this is sufficient for our purposes given that the income data are in a similar form. We plot house prices (Y_{tq}) against gross household income (X_{tq}) for each year from 1999 to 2006. The results are shown in Figure 3.

Insert Figure 3 Here

The quantile plots for every year again reveal a strong linear relationship between house prices and income. The results of cross-section regressions of house price deciles against income deciles as described in (1) are shown in Table 3. The R^2 coefficients are even higher than for Sydney, in every case approximating 1.00 to two decimal places.

Insert Table 3 Here

Using the same data sources it is also possible to compare income and house price quantiles at the state level (i.e., for Texas itself) from 2000 to 2006. The quantile plot and corresponding regression equations are provided in Figure 4 and Table 4 respectively. For Texas as a whole again the relationship is clearly linear with an R^2 coefficient approximating 1.00 to two decimal places.

Insert Figure 4 Here

Insert Table 4 Here

In the remainder of this section, we try to shed some light on these remarkable results. Our starting point is a variant on the permanent income hypothesis. We assume that the amount Y_{ti} a household i purchasing a house in period t is willing to spend depends on the deposit paid (D_{ti}) and gross permanent income (X_{ti}^P) from all sources (e.g., labor and financial assets) as follows:

$$Y_{ti} = D_{ti} + \rho_t \sum_{n=0}^{N_t} \left[\frac{X_{ti}^P}{(1+r_t)^n} \right], \tag{2}$$

where

$$X_{ti}^P = r_t \sum_{m=0}^{\infty} \left[\frac{E_t X_{t+m,i}}{(1+r_t)^m} \right],$$

ρ_t is the proportion of gross permanent income that the household is willing to allocate to mortgage repayments, N_t is the length of the mortgage and r_t is the mortgage interest rate. We assume here for simplicity that mortgage repayments per period are fixed and that ρ_t and N_t are the same for all households in period t. The parameters ρ_t and N_t, however, can evolve over time in response to changes in preferences and conditions in the housing market. Empirical support for the assumption that ρ_t does not vary much with income is provided by Piazzesi, Schneider and Tuzel (2007). Using data from the consumer expenditure survey for the years 1984 to 2002 they find that the lowest income quintile on average spends 17.8 percent of gross income on housing while the highest income quintile spends 16.9 percent.

Given that we do not have data on deposits paid, we now further assume that deposits are proportional to house prices. It follows that $D_{ti} = \delta_t Y_{ti}$, where δ_t is typically set by the mortgage market. Substituting for D_{ti} in (2) and using the fact that permanent income X_{ti}^P does not depend on n and hence can be taken outside the summation sign in (2), we obtain that

$$Y_{ti} = \left[\frac{\rho_t}{(1-\delta_t)} \right] \left[\frac{1-(1+r_t)^{-N_t}}{r_t} \right] X_{ti}^P = b_t^* X_{ti}^P. \tag{3}$$

We have implicitly assumed here that financial wealth (W_{ti}) satisfies the following constraint: $W_{ti} \geq \delta_t b_t^* X_{ti}^P$. In other words, when a household's financial wealth is initially below this level, purchase is delayed until this inequality is satisfied. When a household's financial wealth is above this level, the residual is invested elsewhere.

Equation (3) implies that a linear relationship exists between house prices and permanent income. Increases in ρ_t (the proportion of gross permanent income devoted to mortgage interest payments), δ_t (the downpayment ratio), and N_t (the term of the mortgage) and falls in r_t (the cost of borrowing) will act to increase house prices.[3] Given estimates of ρ_t, r_t N_t and δ_t it is possible to estimate b_t^*. We attempt this for Sydney and Houston. Estimates of ρ_t for Sydney are provided by the reciprocal of the REIA/AMP index multiplied by 10.[4] The REIA/AMP index has a value of 0.324 in March 1996, falling to 0.267 in March 1998 before rising gradually to 0.372 in December 2006. There is evidence that average loan length N_t has also risen in recent years. According to Bourassa (1996), $N = 20$ in 1989/1990 in Australia. By 2004, according to the OECD (2004), typical mortgage loan terms in Australia had risen to $N = 25$. Brischetto and Rosewall (2007) report a further increase of loan terms to $N = 30$. We will assume that N rises at a constant rate between 1990 and 2006 from 20 to 30. Data on the mortgage interest rate r_t over the period 1996 to 2006 are obtained from the Reserve Bank of Australia Indicator Lending Rates Table. OECD (2004) estimates that the down-payment ratio δ_t equals 0.35 for Australia.[5]

The mortgage length N_t for Houston is set to 30. The 30-year fixed rate mortgage interest rate r_t is provided by HSH Associates' National Mortgage Statistics. Following Green and Wachter (2004), we set the down-payment ratio δ_t equal to 0.25. These estimates correspond reasonably well with those used by McCarthy and Peach (2004) who assume that households spend a maximum of 27 percent of gross income on mortgage repayments (i.e., $\rho_t = 0.27$), that they hold 30 year loans (i.e., $N = 30$), and that the down-payment ratio δ_t is 0.2. Empirically, they find it fluctuated between 0.23 and

[3]The interpretation of changes in δ_t can be problematic. On the one hand, greater financial wealth exerts upward pressure on house prices. On the other hand, a decline in the required down-payment ratio may encourage previously credit constrained households to enter the housing market earlier or at a higher level than they otherwise would have.

[4]The REIA/AMP index is defined as the ratio of median family gross income to average loan repayments [see Real Estate Institute of Australia and AMP Banking (2007).]

[5]Purchasers of a loan with a value of δ_t below 0.2 in Australia are required to pay Lenders Mortgage Insurance (LMI) – this is known as Private Mortgage Insurance (PMI) in the US. In the event of default on the loan, LMI protects the lender not the borrower. LMI creates an additional significant transaction cost for low δ_t purchasers.

0.26 between 1984 and 2003 in the US. By comparison, according to the OECD (2004), $\delta_t = 0.22$ in the US.

Plugging these numbers into (3), we obtain estimates of b_t^* for Sydney in Table 5 and Houston in Table 6. A comparison of Table 2 with Table 5 reveals that $\hat{b}_t^* > \hat{\beta}_t$ for Sydney in each of 1996, 2001 and 2006. The same pattern is observed for Houston (see Tables 3 and 6). This finding can be attributed to the fact that the explanatory variable in (1) is observable income, while in (3) it is unobservable permanent income. Replacing permanent income with observable income causes attenuation bias in the estimated slope coefficient [see Friedman (1957) and Meghir (2004)]. That is, supposing it is true that $Y_{ti} = b_t X_{ti}^P$, then it follows that when estimating (1) that

$$\text{plim}(\hat{\beta}_t) = \left[\frac{\text{var}(X_{ti}^P)}{\text{var}(X_{ti}^P) + \text{var}(X_{ti}^T)}\right] b_t < b_t,$$

where X_{ti}^T denotes transitory income. Attenuation bias also explains the positive intercepts in (1) observed for Sydney, Houston and Texas.

Insert Table 5 Here

Insert Table 6 Here

The higher values of $\hat{\beta}_t$ and $\hat{b}_t^*$ in Sydney compared with Houston and its faster rate of increase over time can probably be attributed to very high levels of immigration in Sydney [see Robertson (2006)], tough zoning restrictions [again see Robertson (2006)] and the geographical constraints provided by the ocean on one side and the blue mountains on the other.[6] These factors help explain why Sydney was ranked as the seventh most unaffordable market (after five coastal markets in California and Honolulu) in the third annual Demographia International Housing Affordability Survey published in 2007, while by contrast Houston was tied in 122nd place.

[6]In the US context, Glaeser and Gyourko (2003) have also emphasized the importance of zoning restrictions for the evolution of house prices.

3 Some Implications of Our Findings

3.1 Changes in Average House Prices

One common measure of the change in average house prices is a median index.[7] Using (3), the change in a median index can be decomposed into two parts:

$$\ln\left(\frac{Y_{t+1}^{med}}{Y_t^{med}}\right) = \ln\left(\frac{b_{t+1}^*}{b_t^*}\right) + \ln\left(\frac{X_{t+1}^{med,P}}{X_t^{med,P}}\right). \tag{4}$$

The first part $\ln(b_{t+1}^*/b_t^*)$ describes the contribution of changes in the mortgage market, while the second part $\ln(X_{t+1}^{med,P}/X_t^{med,P})$ represents the change in median permanent income over time.[8,9] This decomposition can be useful for policy purposes as a means for understanding trends in the housing market. The two driving forces – income and the mortgage market – combine multiplicatively to determine the evolution of median house prices. It follows that, in the event of government intervention in response to concerns about sustainability and declining affordability, the focus should be on trying to manage the path of the parameter b_t^*.

[7]The interpretation of such an index can be problematic when the mix of houses sold varies from one period to the next. For this reason, hedonic or repeat sales methods should be used to construct quality adjusted indexes [see Diewert (2007)]. Nevertheless, a median index has the advantage that it is easy to compute and is widely available.

[8]This decomposition only requires (3) to hold for the median. Hence we do not require the assumptions that ρ_t and δ_t are the same for all households. All we need are the median's values of ρ_t and δ_t each year.

[9]This equation is not easily tested empirically for two reasons. First, median permanent income is not directly observable. Second, ρ_t in the US is typically calculated as the repayment on the median price house divided by median income while in Australia it is calculated as the average mortgage repayment divided by median income. Either way, this introduces some circularity into (4) when a numerical estimate of ρ is inserted and median permanent income is replaced by observable median income.

3.2 Changes in the Price-to-Income Ratio and the Efficiency of the Housing Market

A large literature has investigated the efficiency of the housing market. Case and Shiller (1989) find evidence of persistence in year on year price changes. The persistence is strong enough to create intertemporal arbitrage opportunities even after allowing for transaction costs. Hill, Sirmans and Knight (1999) also reject the random walk hypothesis for house prices. Given the cross-section nature of our analysis, we are not able to directly address this issue. To the extent that there may be persistence in changes in permanent income and b_t^* our results need not be inconsistent with their findings.

A second strand of the literature explores the existence of price bubbles in the housing market. There is a tendency to equate variations in the price-to-income ratio from its long-run average to departures from equilibrium [see for example Malpezzi (1999), Capozza, Hendershott, Mack and Mayer (2002), and Woodall (2003)]. By contrast, McCarthy and Peach (2004) and Himmelberg, Mayer and Sinai (2005) both argue that the rapid rise in price-to-income ratios in the US since 1995 can be attributed to the concurrent fall in mortgage interest rates, demographic factors, supply constraints and increased competition in the mortgage market. Case and Shiller (2003) and Haines and Rosen (2007) are more nuanced in their conclusions. They find evidence of overheating in some sectors of the US market. Case and Shiller (2003) differ from the others in that they also make use of survey data in their analysis. The evidence for the UK market is also mixed. Cameron, Muellbauer and Murphy (2006) find no evidence of a bubble while in contrast Garino and Sarno (2004) do.

Our results can help shed light on this issue. From (3), it follows that $Y_t^{med}/X_t^{med,P} = b_t^*$.[10] There is no particular reason to expect b_t^* to be stationary. The equilibrium value of b_t^* is likely to change over time in response to deregulation of financial markets, demographic shifts and changes in zoning laws. It follows that we should not expect house

[10]Again, this result does not require ρ_t and δ_t to be the same for all households.

prices and income to be cointegrated. Empirical support for this conclusion is provided by Gallin (2006).

In our context, a debate over the existence of bubbles reduces to a discussion of whether or not b_t^* has reached an unsustainable level. This may happen if ρ_t rises to a level that can only be justified by the expectation of capital gains. It is difficult to say when such a point has been reached. For example, immigration combined with supply constraints can push ρ_t to a level that might seem unsustainable – and indeed would be in the absence of these forces – without actually being so. Certainly a value of $\rho_t = 0.37$ for Sydney in 2006 is high. Whether or not it is sustainable remains to be seen.

3.3 Hedonic Models

One important implication of our results in Tables 2, 3 and 4 is that housing characteristics only affect house prices to the extent that they affect a house's relative position in the overall ordinal ranking of houses (assuming preferences are sufficiently homogeneous across the population to allow the construction of such a ranking). For example, if a house acquires an extra 100 square meters of land, it rises in price because it moves up in the ordinal ranking of houses. The actual increase in price is then determined from the income distribution.

This argument is problematic for parametric hedonic models, especially linear models, which attempt to compute shadow prices for characteristics such as the number of bedrooms, number of bathrooms, land area, location, etc. [see for example Witte, Sumka and Erekson (1979), Linneman (1980), Malpezzi, Chun and Green (1998), Crone, Nakamura and Voith (2004), Stevenson (2004), Gouriéroux, C. and A. Laferrère (2006) and Hill and Melser (2008)]. Our model implies that the concept of a shadow price may not be particularly meaningful in this context since the impact on price of an improvement in a characteristic may be highly sensitive to the initial position of a house in the ordinal ranking, and the shape of the income distribution. This finding argues in favor of semiparametric or nonparametric methods which impose few or no restrictions on functional form [see for example Pace (1993), Gelfand, Ecker, Knight and Sirmans

(2004) and Parmeter, Henderson and Kumbhakar (2007)].

3.4 Policy Implications

Our analysis highlights the importance of understanding the forces driving changes in house prices. We have shown how the parameters ρ_t, δ_t, r_t, N_t in the mortgage market, can determine the house price distribution for a given income distribution. The overall relationship between observed income and house prices is essentially captured by just two (admittedly time dependent) parameters, the intercept and slope, in our linear model. The simplicity of this result should facilitate more informed discussion of the problem of housing affordability and the sustainability of house price movements.

House prices can rise either due to a movement along the income-house price line or due to a shift in the line. Rises that are attributable to the latter are of potentially greater concern to policy makers. The slope b_t^* of the line may rise as a result of a fall in the mortgage interest rate r_t, or increases in ρ_t or N_t. The problem with a rise in b_t^* is that it implies that households are becoming more highly leveraged. If at a later date the mortgage interest rate rises or lending restrictions are tightened, this can impose stress on these households, as has happened recently in the subprime market in the US. The contribution of changes in r_t, ρ_t and N_t to the evolution of house prices in Sydney can be seen in Tables 2 and 6. Between 1996 and 2006, $\hat{\beta}_t$ rose by 62 percent and $\hat{b}_t^*$ by 49 percent. Over this same period, the median house price in Sydney rose by 136 percent [see Australian Property Monitors (2006)].

4 Conclusion

It has been shown that a linear relationship exists between income and house price quantiles for Sydney, Houston and Texas. Such is the strength of the relationship that we suspect that it may hold more generally for other cities and states. We have shown how such a linear relationship can arise from a variant on the permanent income hypothesis. If our finding proves to be robust, it has some interesting implications for our understanding of the housing market. Most significantly, it implies that the house price distribution is more or less independent of the quality of the housing stock itself. Rather, the house price distribution seems to be determined directly from the income distribution via a location-scale transformation which depends on conditions in the mortgage market, and other factors such as demographics and supply constraints. This finding should help inform future discussion on the evolution of house prices.

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Table 1: Summary Statistics for the House Price Distribution (Sydney)

	1996	2001	2006
Mean	261,010	424,470	590,570
Median	218,000	345,000	455,000
Standard Dev.	165,160	311,380	481,270
Skewness	2.613	3.193	3.605
Kurtosis	12.841	17.893	20.895
P-value of Jarque-Bera test	<0.001	<0.001	<0.001
P-value of Jarque-Bera test on log prices	<0.001	<0.001	<0.001

Table 2: Regression of House Price Quantiles against Income Quantiles (Sydney)

Year		α	β	R^2
1996	Coefficient	73758.84	3.09	0.99
	Standard error	4300.71	0.07	
	t-statistic	17.15	44.22	
	Prob $> \|t\|$	0.00	0.00	
2001	Coefficient	87796.58	4.30	0.99
	Standard error	11575.04	0.15	
	t-statistic	7.58	28.93	
	Prob $> \|t\|$	0.00	0.00	
2006	Coefficient	105397.10	5.01	0.95
	Standard error	39196.69	0.42	
	t-statistic	2.69	11.91	
	Prob $> \|t\|$	0.03	0.00	

Table 3: Regression of House Price Quantiles against Income Quantiles (Houston)

Year		α	β	R^2
1999	Coefficient	32229.04	1.77	1.00
	Standard error	1324.78	0.02	
	t-statistic	24.33	79.45	
	Prob > $\|t\|$	0.00	0.00	
2000	Coefficient	30066.97	2.11	1.00
	Standard error	1185.34	0.02	
	t-statistic	25.37	103.24	
	Prob > $\|t\|$	0.00	0.00	
2001	Coefficient	36879.72	1.97	1.00
	Standard error	1217.77	0.02	
	t-statistic	30.28	100.39	
	Prob > $\|t\|$	0.00	0.00	
2002	Coefficient	49827.94	1.87	1.00
	Standard error	1470.16	0.02	
	t-statistic	33.89	81.61	
	Prob > $\|t\|$	0.00	0.00	
2003	Coefficient	55850.78	1.87	1.00
	Standard error	2160.40	0.03	
	t-statistic	25.85	54.16	
	Prob > $\|t\|$	0.00	0.00	
2004	Coefficient	51664.68	1.98	1.00
	Standard error	1874.74	0.03	
	t-statistic	27.56	66.69	
	Prob > $\|t\|$	0.00	0.00	
2005	Coefficient	59008.29	1.99	1.00
	Standard error	1835.20	0.03	
	t-statistic	32.15	70.13	
	Prob > $\|t\|$	0.00	0.00	
2006	Coefficient	58381.61	2.02	1.00
	Standard error	2187.43	0.03	
	t-statistic	26.69	63.53	
	Prob > $\|t\|$	0.00	0.00	

Table 4: Regression of House Price Quantiles against Income Quantiles (Texas)

Year		α	β	R^2
2000	Coefficient	28940.83	1.79	1.00
	Standard error	1239.21	0.02	
	t-statistic	23.35	95.08	
	Prob > $\|t\|$	0.00	0.00	
2001	Coefficient	35025.07	1.73	1.00
	Standard error	843.96	0.01	
	t-statistic	41.50	138.28	
	Prob > $\|t\|$	0.00	0.00	
2002	Coefficient	38868.87	1.71	1.00
	Standard error	1429.25	0.02	
	t-statistic	27.20	83.12	
	Prob > $\|t\|$	0.00	0.00	
2003	Coefficient	41032.53	1.74	1.00
	Standard error	1591.16	0.02	
	t-statistic	25.79	75.51	
	Prob > $\|t\|$	0.00	0.00	
2004	Coefficient	41453.49	1.73	1.00
	Standard error	1990.75	0.03	
	t-statistic	20.82	62.29	
	Prob > $\|t\|$	0.00	0.00	
2005	Coefficient	44639.69	1.80	1.00
	Standard error	2679.16	0.04	
	t-statistic	16.66	48.56	
	Prob > $\|t\|$	0.00	0.00	
2006	Coefficient	43504.56	1.85	1.00
	Standard error	4629.72	0.06	
	t-statistic	9.40	30.62	
	Prob > $\|t\|$	0.00	0.00	

Table 5: Trends in Mortgage Conditions and the Price-to-Income Ratio (Sydney)

	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006
b_t^*	4.43	4.94	5.14	5.49	5.17	5.57	5.95	6.52	7.11	6.81	6.62
P_t^{med} in 1000s of AUS$	218	243	265	292	310	345	399	442	460	458	455
$\bar{X}_t$ in 1000s of AUS$	58.1	63.0	61.9	64.5	67.0	71.0	74.4	78.1	79.2	81.7	83.4
P/X ratio	3.75	3.86	4.28	4.52	4.62	4.85	5.35	5.65	5.80	5.61	5.45

Table 6: Trends in Mortgage Conditions and the Price-to-Income Ratio (Houston)

	1999	2000	2001	2002	2003	2004	2005	2006
b_t^*	2.11	2.12	2.20	2.26	2.36	2.43	2.52	2.57
P_t^{med} in 1000s of US$	101	114	122	130	133	134	142	149
$\bar{X}_t$ in 1000s of US$	56.1	56.0	59.6	60.4	60.9	61.7	64.0	69.4
P/X ratio	1.80	2.04	2.05	2.15	2.24	2.17	2.22	2.14

Figure 1. The House Price Distribution for Sydney in 1996, 2001 and 2006

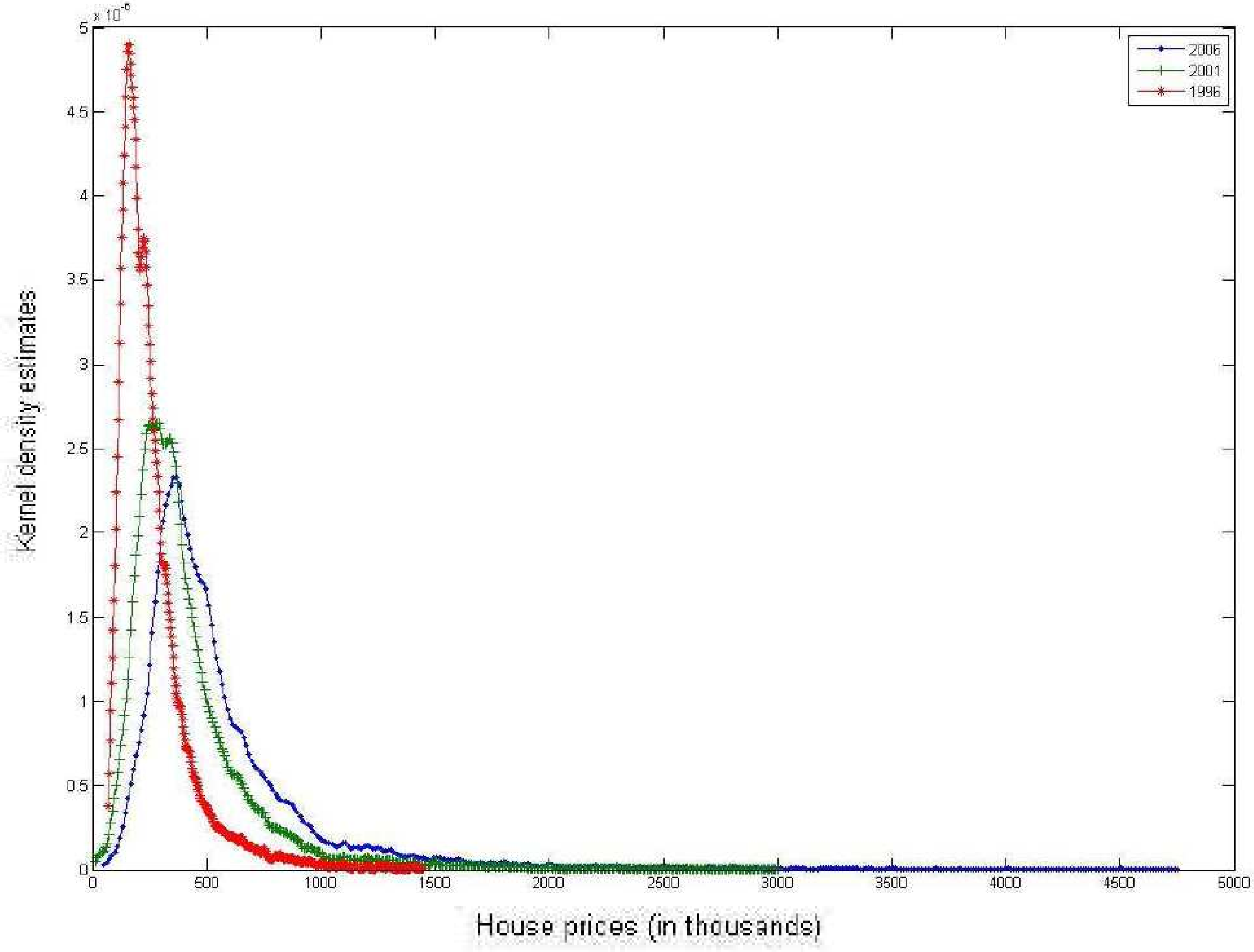

Figure 2. Decile-Decile Plot and Fitted Regression Line of Gross Household Income Against House Prices for Sydney

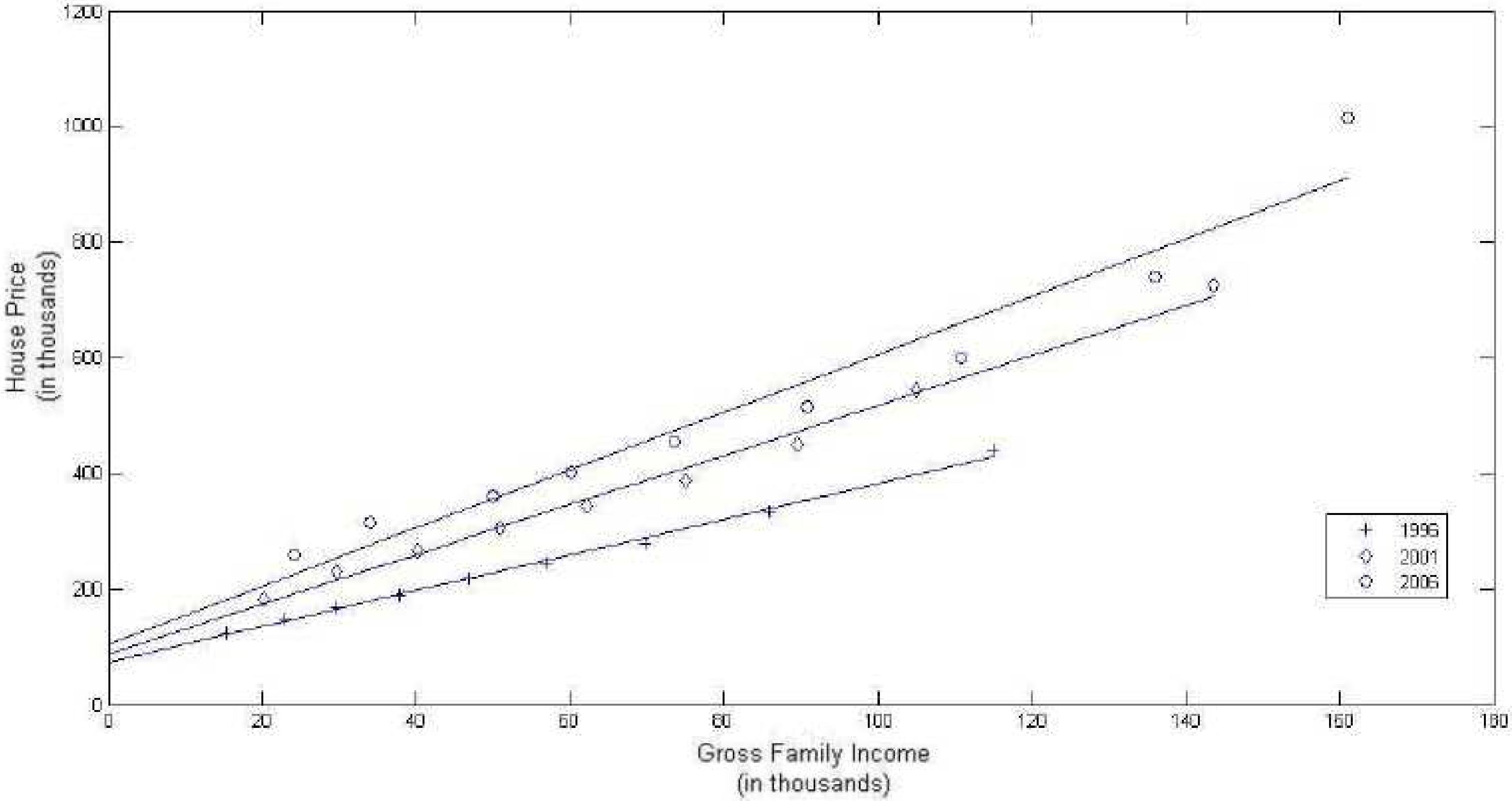

Figure 3. Decile-Decile Plot and Fitted Regression Line of Gross Household Income Against House Prices for Houston

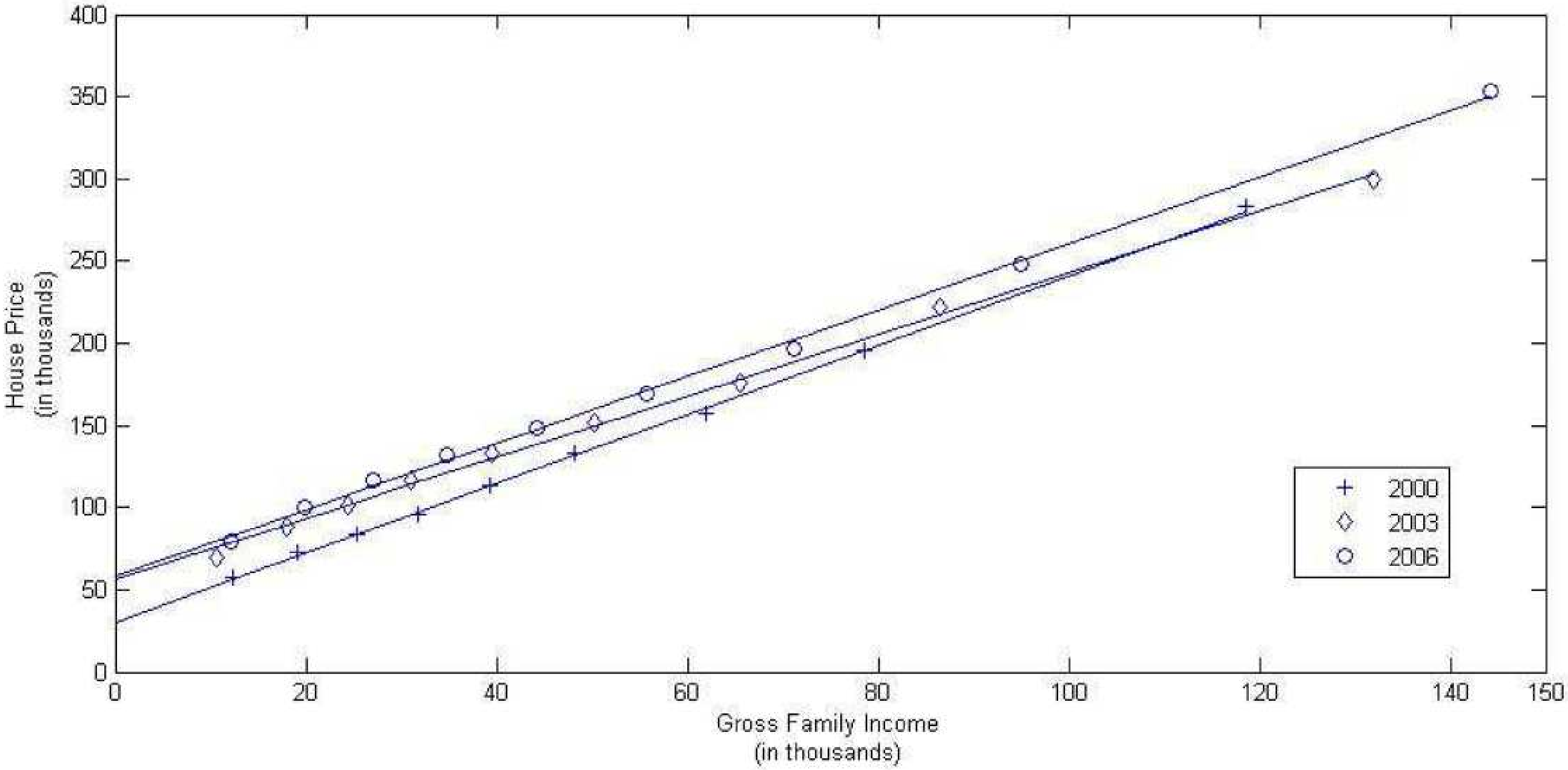

Figure 4. Decile-Decile Plot and Fitted Regression Line of Gross Household Income Against House Prices for Texas