A Re-evaluation of Housing Wealth Effect in Korea*

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Abstract

This paper attempts to estimate housing wealth effect in Korea. As aggregate consumption data include renters' consumption as well as home-owners' consumption, the correlation between aggreagte consumption and housing wealth stock may underestimate the genuine housing wealth effect. We argue that non-housing consumption is more appropriate measure of consumption, because housing consumption is in large part the imputed rents which home- owenrs do not actually pay. Furthermore, we note that consumption share of home-owners needs to be taken into account. Using structural models, we estimate the share of home owners' consumption to be about 65%, suggesting that the magnitude of housing wealth effect in Korea is likely to be larger than what simple time series regressions imply.

Keywords: consumption, housing wealth, home ownership

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1 Introduction

The last decade witnessed dramatic changes in housing prices in many countries. According to Sutton (2002), housing price in the U.S. increased by 21% net of inflation during 1995-2001, and during the same period the real rates of increases in housing prices reached 42%, 60%, and 70% in the U.K., Netherlands, and Ireland, respectively. Such steep rises in housing prices are naturally expected to affect aggregate demand and economic activities via various channels, one of which is wealth effect. In fact, quite a few media articles have emphasized the strong private consumption via housing wealth effect as a factor that helped the U.S. economy weather the 2001 recession smoothly.

In a standard life cycle or permanent income model, higher housing wealth should lead to increases in consumption even when there are bequest motives or borrowing constraints. In contrast to such clear theoretical predictions, empirical evidence for the magnitude of housing wealth effect is mixed: For example, Case, Quigley, and Shiller (2005) report that the elasticity of consumption with respect to housing wealth ranges from 0.1 to 0.17 in the panel analyses of 14 major countries, but Girouard and Blondal (2001) estimate the elasticities for the U.S., U.K., and France to be 0.02, 0.06, and 0.08, respectively, which are much lower than in Case, Quigley, and Shiller.

In this paper, we attempt to estimate the magnitude of housing wealth effects in the case of Korea. In doing so, we emphasize two factors, i.e., home ownership rate and the proper measure of consumption, that may resolve the discrepancy between the theory and empirics in housing wealth effect. An intuitive story regarding home ownership goes as follows: home-owners may well perceive house price increases as an addition to their wealth and therefore increase their consumption given current income, while renters planning to purchase their own homes and home-owners wanting to trade up may decrease their consumption as they will have to save more for higher down payments and repayments. Therefore, if estimated at the aggregate (across home-owners and renters) level, the magnitude of housing wealth effect will be inversely related with the fraction of renters, or more appropriately, the weight of renters' consumption. We also note that housing consumption, a sizable component of total consumption, is proxied by imputed rents which are the estimate of what would be paid if the owner-occupied dwellings were rented. Seeing that higher housing prices generally accompany higher rent payments, using total consumption will lead to underestimation of the housing wealth effect.

To verify these insights, we first run time series regressions with aggregate (across home-owners and renters) data and estimate the elasticity of consumption with respect to gross housing value. Regression results support the presence of significant housing wealth effect on non-housing consumption, while higher housing value tends to lower housing consumption significantly. It is also shown that the negative effect on housing consumption is large enough to render housing wealth effect on total consumption (i.e., the sum of housing and non-housing consumption) insignificant or of the opposite sign. We interpret these findings as consistent with the claim that it is more appropriate to exclude housing consumption from the total consumption when using time series data.

We then proceed to run cross section regressions with data on home-owners only, and we find that the cross sectional estimates of consumption elasticity are much higher than those obtained from time series regressions.

Although caution should be exercised in comparing the two sets of regression results, we view the considerable difference between the results is mainly due to the fact that a sizable fraction of households in Korea are renters.[1] Intuitively, after the rise in housing rents that is usually accompanied by higher housing prices, renters are left with less disposable resources. Unless the elasticity of substitution between housing and non-housing consumption is large, higher housing price will decrease the non-housing consumption of renters let alone their housing consumption. Hence, the higher the proportion of renters is, the lower the aggregate time series estimates of housing wealth effect are compared to what would result from the regressions with home-owners only.

We next turn to the question of how much to revise the aggregate time series measures of wealth effect, taking the home ownership heterogeneity into account. To do so, we set up simple stochastic dynamic models populated by home-owners and renters, and estimate the consumption shares of each consumer group. We find that the revised time series estimates of housing wealth effect is of similar magnitude to the cross sectional estimates.

This paper is organized as follows. In Section 2, we estimate housing wealth effect in Korea with aggregate time series and cross section data and consider if the results are consistent with our claims above. Section 3 is devoted to the estimation of home-owners' consumption share in the context of dynamic structural models, and to a possible revision of housing wealth effect estimated using time series data.

[1] The proportion of home-owners in Korea is estimated to be around 52.8% on average during 1995-2000 and 61% in 2001, which are lower than in the US (67.5%) and UK (67.0%) reported in Ludwig an Slok (2002).

2 Measuring Housing Wealth Effect

2.1 Time Series Results

To obtain benchmark estimates of housing wealth effect with aggregate time series data, we focus on long-run relationships rather than short-run correlations among consumption, income, and wealth. This estimation strategy stems from the view that wealth effect is intrinsically a long-run concept: a representative consumer's inter-temporal budget constraint dictates that, when his wealth rises, he must spend it over lifetime unless he transfers all of the money to others. Considering the possibility that consumption responds to changes in asset value with a substantial lag or over a long time period, it is more appropriate to estimate wealth effect in the long-run.

From the perspectives of time series econometrics, long-run relationships are estimated by cointegrating regressions: for example, building on Campbell and Mankiw (1989), Lettau and Ludvigson (2000) derive from a representative consumer's intertemporal budget constraint a long-run relationship among consumption, labor income and asset wealth, and show that they should be cointegrated.

Using the long-run relationship of the Korean quarterly data for the sample period 1987:I to 2003:IV, we estimate the following equation:

$$C_t^i = \beta_0^i + \beta_y^i Y_t + \beta_h^i HW_t + \beta_s^i SW_t + \epsilon_t \quad (1)$$

where $i = 1, 2$, and 3 denotes total consumption, non-housing consumption, and housing consumption, respectively. The housing consumption (C^3) is proxied by imputed rents, and the non-housing consumption (C^2) is calculated as the residual of total consumption net of the housing consumption. Y_t denotes real Gross National Income. The income and consumption variables are all seasonally adjusted series. Housing wealth (HW) is constructed as the product of nationwide housing price index and the linearly interpolated annual series on the number of dwellings. We divide nominal housing wealth series by CPI to obtain real series. Stock wealth (SW) is proxied by the value of stock held by individuals, available from the flow of funds table, deflated by CPI. Each variable is transformed into per capita terms in logs.[2]

Table 1 reports the estimation results for the three consumption series. The first panel shows that the estimated elasticities of the total consumption with respect to both the housing and stock wealth are either of the wrong signs or statistically insignificant. Including a time trend in the regression do not alter the results qualitatively. In contrast, as can be seen in the second panel, the estimation results for non-housing consumption support significant housing wealth effect. The elasticity of non-housing consumption with respect to housing wealth is estimated to be 0.11 with a time trend term and 0.08 without it, and in both cases the estimates are statistically significant. However, the estimated stock wealth effect again turns out to be insignificant.[3]

Turning to the third panel reporting the results for housing consumption, we find a significantly negative correlation between housing consumption and housing wealth, suggesting that demand for housing service is likely to decline as the housing price rises. This finding explains why housing wealth effect on the total household consumption turns out to be smaller than on the

[2] ADF tests applied to the data series on (C^i, Y, HW, SW) could not reject the null of unit roots even at 10% significance level, except that the null is not rejected for stock wealth at 1% level. Also, residual based ADF tests support the presence of cointegration relations.

[3] Larger effect of housing wealth on consumption than of stock wealth is also reported in Case, Quigley and Shiller (2005).

non-housing consumption: economically as well as statistically insignificant estimate of housing wealth effect on total consumption is attributable to the adverse effect of housing price on housing consumption.[4]

A closer look into the results in Table 1 and how aggregate housing consumption is constructed suggests that non-housing consumption is a more appropriate variable when evaluating housing wealth effect: imputed rents for dwellings occupied by owners, the main component of housing consumption, are the estimates of the rent that would be paid if the dwellings were rented housing, rather than the actual amount spent by home-owners during a given period of time. Therefore, using total consumption inclusive of housing consumption leads to the under-estimation of the true housing wealth effect, in view of the negative correlation between housing consumption and housing prices.

Even if one concentrates on non-housing consumption, however, there is another reason why one may not get a precise estimate of housing wealth effect from the aggregate time series. It is because the "textbook" definition of housing wealth effect is appropriate for home-owners, but aggregate non-housing consumption is the sum of renters' consumption and home-owners' consumption. As for the home-owners, the rise in housing price will increase non-housing consumption via the favorable substitution and wealth effect. After the rise in housing rental costs usually accompanied by higher property prices, however, renters left with less disposable resources will decrease housing consumption, while at the same time they will substitute non-housing consumption for housing consumption. Therefore, if the elas-

[4] In the National Income Account of Korea, housing consumption takes up about 15% of total consumption expenditure.

ticity of substitution between housing and non-housing consumption is low enough, renters' non-housing consumption can decrease. Even in the opposite case, we can expect that the responsiveness of home-owners' housing consumption will be higher than that for renters. In a nutshell, it is highly likely that the estimated elasticities of aggregate non-housing consumption in Table 1 is smaller than the true measure of wealth effect.

2.2 Cross Section Results

In the previous subsection, we emphasized the possibility that, if aggregate time series data are used, ignoring home ownership is a main reason for underestimating the degree of housing wealth effect. In this section, we use cross sectional data to get another measure of housing wealth effect. In particular, our aim here is to see whether the magnitude of housing wealth effect estimated with cross sectional data on home-owners only is considerably higher than what is estimated by time series data without distinguishing home ownership profile.

We use the following specification

$$C_{j,t} = \beta_0 + \beta_y Y_{j,t} + \beta_h HW_{j,t} + \beta_z X_{j,t} + \epsilon_{j,t} \tag{2}$$

where $(C_{j,t}, Y_{j,t}, HW_{j,t})$ denote total consumption expenditure, household income, and housing wealth, respectively, of the j^{th} household at period t, and X is a vector of controls such as lagged consumption, household size, and education years and ages of household heads. Two variants of equation (2) are considered depending on the inclusion of stock wealth as a control in X, and for each variant we also consider the inclusion/exclusion of lagged consumption. All variables are in logarithmic terms except the household size, education years, and ages. The data series are obtained from Korea Labor and Income Panel Study (KLIPS). Although the KLIPS dataset was constructed for 1998-2001, only two years' observations are available for the specification including stock wealth and lagged consumption. Therefore, we used pooled cross sectional regressions instead of formal panel regressions.

The estimation results are summarized in Table 2. One conspicuous finding is the significance of the coefficients on housing wealth for all specifications: in the first column, for example, a household with 1% higher housing wealth than the other tends to spend more on consumption by 0.16%. When lagged consumption is included in X, this number slightly decreases to 0.11%, but still remains significant as shown in the second column.

Another finding is that the coefficients on housing wealth are almost invariant to the inclusion/exclusion of stock wealth as a control. We attribute this invariance to the very small fraction of household investing in stock or equity, as seen from the number of observations in the last row. Also, the proportion of the total population participating in the equity market is reported to be only 2.9% in 1997 and 8.2% in 2003.

Comparing the results in Table 2 with those in the second panel of Table 1, we observe that the cross-sectional elasticities of home-owners' consumption are larger than the aggregate (over the home ownership profile) time series coefficients. Although direct comparison of the two sets of results is hard to warrant, we interpret this finding as consistent with our conjecture on the importance of ownership profile in measuring housing wealth effects.[5]

One natural suggestion now would be to track down the fraction of homeowners (or, more appropriately, their consumption share) across time and take this information into account when estimating wealth effect using aggregate time series data. Unfortunately, at least in the case of Korea, the

[5] In fact, there is a possibility that the results in Table II may have underestimated housing wealth effect for home owners, because the consumption series used to estimate (2) is the expenditure on total (not non housing) consumption.

only available data of this property are home ownership ratios, which are available only every a few years. In the next section, therefore, we attempt to get an estimate of the consumption share of home-owners, and use the result to re-interpret the time series estimates of housing wealth effect.

3 Estimation of Consumption Shares

In the previous section, we have emphasized the importance of considering home ownership profile when one tries to measure housing wealth effect using aggregate time series. One way to address this issue is to revise the time series estimates of housing wealth effect in view of the relative weight on home-owners's consumption. In this section, therefore, we attempt to estimate the share of home-owners' consumption in the context of dynamic structural models using macro time series.

We consider two simple dynamic models of endowment economy, Models (I) and (II), both populated by home-owners and renters. In Model (I), households maximize lifetime utility defined over non-housing consumption (C_1) and housing consumption (C_2). The sole difference between home-owners and renters in this model is that only the former group has access to the market for housing investment good. In Model (II), renters are further restricted to live by "rule-of-thumb" in the sense of Campbell and Mankiw (1989), living out of their current income period by period.

Our approach here belongs to the literature on the estimation of Euler equation including Campbell and Mankiw (1989), Jappelli and Pagano (1989), and Iacoviello (2004): we first derive optimality conditions for each group of households, combine them into an aggregate Euler equation, and estimate the consumption weight on home-owners.

3.1 Models

In Model (I), the lifetime utility for both types of households to maximize is given by

$$E_0 \sum_{t=0}^{\infty} \beta^t \frac{1}{1-\frac{1}{\sigma}} u(C_{1t}^j, C_{2t}^j)^{1-\frac{1}{\sigma}}, \quad 0<\beta<1, \quad \sigma>0 \tag{3}$$

where the composite consumption $u(C_{1t}, C_{2t})$ is in turn defined as

$$u(C_{1t}, C_{2t}) = \left(\omega^{\frac{1}{\varepsilon}} C_{1t}^{\frac{\varepsilon-1}{\varepsilon}} + (1-\omega)^{\frac{1}{\varepsilon}} C_{2t}^{\frac{\varepsilon-1}{\varepsilon}} \right)^{\frac{\varepsilon}{\varepsilon-1}}, \quad 0<\omega<1, \quad \varepsilon>0 \tag{4}$$

with ε and σ being the elasticity of substitution between the two types of consumption and the (inverse of) elasticity of intertemporal substitution, respectively. The superscript $j = o, r$ denotes home-owners and renters. Note that the form of the composite consumption implies the following relation between the two consumption components (C_1, C_2) and their prices (P_1, P_2) :

$$\left(\frac{\omega}{1-\omega}\right)^{\frac{1}{\varepsilon}} \left(\frac{C_{1t}}{C_{2t}}\right)^{-\frac{1}{\varepsilon}} = \frac{P_{1t}}{P_{2t}} \tag{5}$$

which can be used to estimate the parameters (ω, ε) by a simple regression.[6]

Our strategy is to derive the optimization conditions for each strand of households, combine them to derive the Euler equation for aggregate consumption, and estimate the (steady state) share Λ of home-owner's consumption. In the Appendix I, we show that, for a vector X_t of variables dated t and earlier, the following condition holds in the log-linearized version of Model (I):

$$E_t\left[M_{t+1} X_t\right] = 0 \tag{6}$$

[6] The OLS estimates of (ε, ω) are (0.454, 0.868), with Newey-West standard errors of (0.00, 0.078) and R^2 of 0.440.

where the primitive moment condition M_{t+1} is obtained from

$$\begin{aligned}\Lambda\left[1-\beta(1-\delta)\right]E_t\widehat{P}_{2,t+1} &= \frac{1}{\sigma}\left[\widehat{C}_{1t}-E_t\widehat{C}_{1,t+1}\right] \\ &+\Theta_1\left[E_t\widehat{P}_{1,t+1}-\widehat{P}_{1t}\right] \\ &-\Theta_2\left[E_t\widehat{P}_{2,t+1}-\widehat{P}_{2t}\right] \\ &-\Lambda\left[\widehat{Q}_t-\beta(1-\delta)E_t\widehat{Q}_{t+1}\right] \\ &+(1-\Lambda)\widehat{R}_t\end{aligned} \quad (7)$$

with

$$\begin{aligned}\Theta_1 &= \varepsilon(1-\Xi)(\frac{1}{\varepsilon}-\frac{1}{\sigma})-1 \\ \Theta_2 &= \varepsilon(1-\Xi)(\frac{1}{\varepsilon}-\frac{1}{\sigma})\end{aligned}$$

and Ξ being the ratio $\frac{\omega^{\frac{1}{\varepsilon}}C_1^{\frac{\varepsilon-1}{\varepsilon}}}{\omega^{\frac{1}{\varepsilon}}C_1^{\frac{\varepsilon-1}{\varepsilon}}+(1-\omega)^{\frac{1}{\varepsilon}}C_2^{\frac{\varepsilon-1}{\varepsilon}}}$ evaluated in steady state. In equation (7), Q and R denote real housing price and real gross interest rate, and all variables with "$\widehat{\ }$" are log-deviations from steady sate levels. We use equation (7) as orthogonality conditions to which we apply GMM.[7]

In the second model economy, the descriptions of home-owners are the same as in Model (I), but we further assume that renters do not smooth their consumption path in the face of fluctuations in their period-by-period endowment income Y_t^r. Therefore, they maximize period by period utility

$$\max \quad u(C_{1t}^r, C_{2t}^r) \quad (8)$$

subject to the budget constraint

$$Y_t^r = P_{1t}C_{1t}^r + P_{2t}C_{2t}^r\ . \quad (9)$$

[7] To simplify the estimation procedure, we opt to estimate the pair of (σ,Λ). The values of (ω,ε) are fixed at the OLS estimates in the Appendix. Ξ is fixed at the sample mean of its data counterpart calculated using the aboves estimates. The rate of depreciation δ of housing stock is fixed at 0.007, matching the annual rate estimated for the U.S. by Harding et al. (2004), and β is fixed at the conventional rate of 0.99.

In Appendix II, we summarize how to combine optimality conditions of the two types of households and derive the Euler equation for aggregate non-housing consumption :

$$
\begin{aligned}
\frac{1-\Lambda}{\sigma} E_t \widehat{Y}^r_{t+1} = {} & \Lambda \left[\widehat{Q}_t - \beta(1-\delta) E_t \widehat{Q}_{t+1} \right] \\
& + \Lambda\Theta_1 \left[\widehat{P}_{1t} - E_t \widehat{P}_{1,t+1} \right] - \Lambda\Theta_2 \left[\widehat{P}_{2t} - E_t \widehat{P}_{2,t+1} \right] \\
& + \Lambda \left[1-\beta(1-\delta)\right] E_t \widehat{P}_{2,t+1} - \frac{1}{\sigma} \left[\widehat{C}_{1t} - E_t \widehat{C}_{1,t+1} \right] \\
& + \sigma(1-\Lambda) Y^r_t - \sigma(1-\Lambda) \left(\Theta_3 + (1-\Theta_3)\varepsilon\right) \widehat{P}_{1t} \\
& - \sigma(1-\Lambda)\left(1-\Theta_3\right)(1-\varepsilon) \widehat{P}_{2t} \\
& + \frac{1-\Lambda}{\sigma} \left(\Theta_3 + (1-\Theta_3)\varepsilon\right) E_t \widehat{P}_{1,t+1} \\
& + \frac{1-\Lambda}{\sigma} (1-\Theta_3)(1-\varepsilon) E_t \widehat{P}_{2,t+1}
\end{aligned} \tag{10}
$$

with

$$\Theta_3 = \frac{\omega P_1^{1-\varepsilon}}{\omega P_1^{1-\varepsilon} + (1-\omega) P_2^{1-\varepsilon}}$$

in steady state. We use (10) to construct moment conditions for GMM estimation.[8]

3.2 Estimation results

We use quarterly Korean data for the period 1987:Q1 to 2003:Q4. The aggregate non-housing consumption C_1 is the same series used for the time series regression in section 2. The prices series (P_1, P_2) of non-housing and housing consumption, respectively, are proxied by corresponding implicit deflators available from national income account, and the housing price Q is the nationwide house price index reported by Kookmin Bank. All price variables are deflated by CPI. The real interest rate R is the 3 year corporate bond rate adjusted for ex-post CPI inflation. For the endowment Y_t^r of

[8] We use the same pre-fixed values for $(\omega, \varepsilon, \Xi, \delta)$ as for Model (I). Also, Θ_3 is fixed at the corresponding sample average.

renters in the second model economy, we use the per capita GNI resorting to the assumption that the endowment profile is identical for all households. Since variables in equations (7) and (10) are represented in log-deviations from steady state levels, I use the Hodrick-Prescott filter to construct correspondingly transformed series.

The left panel of Table III reports the estimation results for Model (I). In column (1), the second to fourth lags of $(C_1, C_2, P_1, P_2, Q, R)$ are used as instruments, while lags of C_2 not appearing in the moment condition are dropped out in column (2). The estimated elasticities of intertemporal substitution imply that the utility function is close to the logarithmic function in both columns, but the estimate 0.769 of Λ in column (2) is higher than 0.659 in column (1). Both parameters are estimated sharply, as reflected in their t-values.

The estimation results for Model (II) are reported in the right panel of Table III. In column (3), the second to fourth lags of $(C_1, C_2, P_1, P_2, Q, R, Y)$ are used as instruments, while lags of C_2 not appearing in the moment condition (7) are again dropped out in column (4). The estimates of σ and Λ for Model (II) tend to be lower than those for Model (I): the estimates of σ, ranging around 0.8 to 0.9, are significantly lower than what for Model (I) at the 5% significance level but not at the 1% level. Those of Λ, now ranging from 0.61 to 0.64, are not much different from what for Model (I). We believe the slightly lower estimates of Λ in Model (II) better represent the situation that renters are more likely to behave as "rule-of-thumb" consumers, rather than being able to smooth their consumption over time.[9] That being the

[9] In year 2000, the home owership ratio in Seoul was 53.4%, and the weighted average of home ownership ratios in seven major cities was 56.6% if we use the number of households in each city as weights. The average home ownership rate nationwide is 61.0%. The estimted share Λ is therefore higher than the "mass" of home owners, which appears plausible: the consumption of a representative home owner will be higher than that of a

case, the consumption share of renters latent in the aggregate data series will be better captured by the latter model. It is also worth noting that the estimate of Λ is not very different from a similar estimate in Campbell and Mankiw (1989): they estimate the "mass" of consumers who do not borrow or save to smooth consumption to be in the neighborhood of 0.4. If we view the renters as unable to draw resources from housing wealth to smooth consumption, then our estimates of Λ are comparable to those in Campbell and Mankiw (1989), although our estimates give the consumption share, not the "mass," of constrained households.

3.3 Interpretation

In this subsection, we develop a heuristic idea on how much to revise the measure of housing wealth effect estimated with aggregate (across home-owners and renters) time series. To simplify the argument, we assume the following relations

$$\begin{aligned} C^o &= \alpha HW + \varepsilon^o \qquad (11) \\ C^r &= \beta HW + \varepsilon^r \end{aligned}$$

where C^o and C^r are the consumption of home-owners and renters, respectively, and HW is the total housing wealth. If the two consumption series were available, the OLS estimate of α would correspond to the "textbook" definition of housing wealth effect, while that of β would measure the degree of negative income effect on renters coming from the increases in housing price.

Now suppose that a relation analogous to (11) is estimated for per capita consumption C, which is the weighted sum $\Lambda C^o + (1 - \Lambda)C^r$ of two strands

representative renter.

of households' consumption. If we denote the resulting estimate by $\hat{\gamma}$, it follows that the OLS estimates $(\hat{\alpha}, \hat{\beta}, \hat{\gamma})$ have the following relation:

$$\hat{\alpha} = \frac{\hat{\gamma}}{\Lambda} - \frac{1 - \Lambda}{\Lambda}\hat{\beta} \ . \tag{12}$$

Equation (12) shows that how well the aggregate estimate $\hat{\gamma}$ reflects the hypothetical "home-owners only" estimate $\hat{\alpha}$ depends on $\hat{\beta}$ and Λ. Obviously, if Λ is close to one, $\hat{\gamma}$ is not much different from the degree of genuine housing wealth effect. However, if Λ is substantially smaller than 1, $\hat{\gamma}$ tends to underestimate α.

Another potential source of underestimation is that β is significantly smaller than $\hat{\alpha}$. As discussed in section 2, the time series regression results in Table I strongly support that $\hat{\beta}$ would be significantly lower than $\hat{\alpha}$ or possibly even negative. If we put $\hat{\beta} = 0$, which appears not much of an extreme case, the estimates Λ in Table III imply that the genuine wealth effect coefficients $\hat{\alpha}$ recovered from the second panel of Table II would be around 0.14 on average. It is interesting that this number is not very different from the cross sectional regression results in Table II.

4 Conclusion

In this paper, we measure the effect of housing wealth on consumption in Korea. From time series regressions, we find that housing wealth is positively correlated with non-housing consumption but negatively correlated with housing consumption. These results suggest that it is better to use non-housing consumption in measuring housing wealth effect: the negative effect on housing consumption falls mainly on renters, who do not accumulate housing wealth. Furthermore, since aggregate non-housing consumption includes renters' consumption, the correlation between non-housing consumption and housing wealth stock may underestimate the genuine wealth effect.

We then turn to the question of how much to revise the aggregate time series estimates of housing wealth effect. We construct two structural models and estimate the share of home-owners' consumption in those models' context. It is suggested that, if properly revised in light of the estimated consumption shares of home-owners, time series measures of housing wealth effect are likely to be larger than what simple regressions imply.

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Table I: Time Series Regression

Consump.	**Total**		**Non-Housing**		**Housing**	
Model	**I**	**II**	**I**	**II**	**I**	**II**
Const.	-0.315	0.552	-1.500	-0.920	2.721	4.974
	(-0.83)	(1.22)	(-3.49)	(-1.72)	(5.87)	(13.05)
trend	-	0.002	-	0.001	-	0.005
	-	(3.15)	-	(1.78)	-	(9.68)
Y	0.982	0.859	0.998	0.915	0.906	0.586
	(32.82)	(17.83)	(29.41)	(16.01)	(24.77)	(14.38)
HW	-0.000	0.052	0.078	0.113	-0.376	-0.242
	(-0.01)	(1.52)	(2.18)	(2.80)	(-9.74)	(-8.41)
SW	-0.010	-0.025	-0.026	-0.036	0.066	0.025
	(-0.38)	(-1.03)	(-0.89)	(0.11)	(2.11)	(1.23)
$\overline{R}^2$	0.99	0.99	0.99	0.99	0.99	0.99

Note: Numbers in parentheses are t-values.

Table II: Cross Sectional Regression

	(1)	(2)	(3)	(4)
Const.	1.372	0.722	1.966	1.248
	(2.77)	(1.56)	(15.23)	(10.32)
Y	0.220	0.132	0.176	.126
	(5.60)	(3.50)	(22.54)	(17.09)
HW	0.156	0.110	0.157	0.094
	(4.94)	(3.69)	(19.36)	(12.20)
SW	0.079	0048	-	-
	(4.23)	(2.72)	-	-
Lag. C	-	0.413	-	0.395
	-	7.55	-	(28.29)
H. size	0.137	0.086	0.155	0.096
	(5.35)	(3.51)	(19.36)	(16.42)
Education	0.085	0.039	0.105	0.059
	(4.50)	(2.12)	(18.34)	(10.72)
Age	0.053	0.032	0.059	0.030
	(2.67)	(1.75)	(13.39)	(7.22)
Age^2	-0.000	-0.000	-0.001	-0.000
	(-2.33)	(-1.61)	(-14.59)	(-8.13)
$\overline{R}^2$	0.492	0.573	0.605	0.670
No. Obs.	303	303	4066	4061

Note: Numbers in parentheses are t-values.

Table III: GMM Estimation of the Euler Equation

	Model I		Model II	
	(1)	(2)	(3)	(4)
σ	1.230	1.267	0.894	0.808
	(3.72)	(3.89)	(18.81)	(13.35)
Λ	0.659	0.769	0.613	0.639
	(2.91)	(3.02)	(2.97)	(2.19)
	-	-	-	-
$Instru.$	$C_{1,t-2},...,C_{1,t-4}$ $P_{1,t-2},...,P_{1,t-4}$ $Q_{t-2},...,Q_{t-4}$ $R_{t-2},...,R_{t-4}$ $C_{2,t-2},...,C_{2,t-4}$	$C_{1,t-2},...,C_{1,t-4}$ $P_{1,t-2},...,P_{1,t-4}$ $Q_{t-2},...,Q_{t-4}$ $R_{t-2},...,R_{t-4}$	$C_{1,t-2},...,C_{1,t-4}$ $P_{1,t-2},...,P_{1,t-4}$ $Q_{t-2},...,Q_{t-4}$ $R_{t-2},...,R_{t-4}$ $Y^r_{t-2},...,Y^r_{t-4}$ $C_{2,t-2},...,C_{2,t-4}$	$C_{1,t-2},...,C_{1,t-4}$ $P_{1,t-2},...,P_{1,t-4}$ $Q_{t-2},...,Q_{t-4}$ $R_{t-2},...,R_{t-4}$ $Y^r_{t-2},...,Y^r_{t-4}$
$J-stat.$	0.764	0.891	0.753	0.778

Note: Numbers in parentheses are t-values calculated by HAC standard errors of New-West (1987). Numbers in the last row are p-values associated with Hansen's (1982) J-test for the model's overidentifying restrictions.

6 Appendix I : Derivation of Equation (7)

We consider a simple discrete time infinite horizon endowment economy, populated by home-owners and renters. Both types of households are assumed to be identical, except that the only the former group can accumulate housing stock.

6.1 Home-owners

A representative home owner maximizes a standard lifetime utility given by

$$\max E_0 \sum_{t=0}^{\infty} \beta^t \frac{1}{1-\frac{1}{\sigma}} u(C_{1t}, C_{2t})^{1-\frac{1}{\sigma}}, \quad 0 < \beta < 1, \quad \sigma > 0$$

where the instantaneous utility $u(C_{1t}, C_{2t})$ is an CES aggregator of non-housing consumption C_{1t} and housing service consumption C_{2t}:

$$u(C_{1t}, C_{2t}) = \left[\omega^{\frac{1}{\varepsilon}} C_{1t}^{\frac{\varepsilon-1}{\varepsilon}} + (1-\omega)^{\frac{1}{\varepsilon}} C_{2t}^{\frac{\varepsilon-1}{\varepsilon}} \right]^{\frac{\varepsilon}{\varepsilon-1}}, \quad 0 < \omega < 1, \quad \varepsilon > 0.$$

We assume the representative home owner receives a random real endowment Y_t priced at P_{1t}, lends (or borrow) B_t and receives the gross interest payment $R_t B_t$ next period. He also purchases I_t^h units of housing stock for investment (priced at Q_t), and the housing stock evolves according to $H_{t+1} = I_t^h + (1-\delta)H_t$, where δ is the rate of depreciation for the existing housing stock. He leases H_t units of housing stock at the price of Z_t to a housing service firm, which in turn produces and provides housing service C_{2t} at the price of P_{2t}. Therefore, the budget constraint of a owner-occupier is

$$Y_t + Z_t H_t + R_{t-1} B_{t-1} = P_{1t} C_{1t} + P_{2t} C_{2t} + Q_t (H_{t+1} - (1-\delta) H_t) + B_t$$

where all price and quantity variables are in real terms.

The first order conditions for (C_1, C_2, H) are given by

$$\omega^{\frac{1}{\varepsilon}} \left[u_t\right]^{\frac{1}{\varepsilon}-\frac{1}{\sigma}} C_{1t}^{-\frac{1}{\varepsilon}} = P_{1t}\mu_t \tag{A1}$$

$$(1-\omega)^{\frac{1}{\varepsilon}} \left[u_t\right]^{\frac{1}{\varepsilon}-\frac{1}{\sigma}} C_{2t}^{-\frac{1}{\varepsilon}} = P_{2t}\mu_t \tag{A2}$$

$$Q_t\mu_t = \beta E_t \mu_{t+1} Q_{t+1}(1-\delta) + \beta E_t \mu_{t+1} Z_{t+1} \tag{A3}$$

It is worth noting equations (A1) and (A2) implies

$$\left(\frac{\omega}{1-\omega}\right)^{\frac{1}{\varepsilon}} \left(\frac{C_{1t}}{C_{2t}}\right)^{-\frac{1}{\varepsilon}} = \frac{P_{1t}}{P_{2t}} \tag{A4}$$

which can be used to pin down the elasticity of substitution between the two kinds of consumption.

After re-arranging log-linearized versions of (A1), (A3), and (A4), we get[10]

$$\begin{aligned} & \hat{Q}_t + \Theta_1 \hat{P}_{1t} - \frac{1}{\sigma}\hat{C}_{1t} - \Theta_2 \hat{P}_{2t} \\ = \; & \Theta_1 E_t \hat{P}_{1,t+1} - \frac{1}{\sigma} E_t \hat{C}_{1,t+1} - \Theta_2 E_t \hat{P}_{2,t+1} \\ & + \beta(1-\delta) E_t \hat{Q}_{t+1} - (1-\beta(1-\delta)) E_t \hat{P}_{2,t+1} \end{aligned} \tag{A5}$$

where

$$\begin{aligned} \Theta_1 &= \varepsilon(1-\Xi)\left(\frac{1}{\varepsilon}-\frac{1}{\sigma}\right) - 1 \\ \Theta_2 &= \varepsilon(1-\Xi)\left(\frac{1}{\varepsilon}-\frac{1}{\sigma}\right) \end{aligned}$$

and Ξ is the ratio $\frac{\omega^{\frac{1}{\varepsilon}} C_1^{\frac{\varepsilon-1}{\varepsilon}}}{\omega^{\frac{1}{\varepsilon}} C_1^{\frac{\varepsilon-1}{\varepsilon}} + (1-\omega)^{\frac{1}{\varepsilon}} C_2^{\frac{\varepsilon-1}{\varepsilon}}}$ evaluated in a steady state.

[10] In deriving (A5), we resort to the assumption that the technology of the competitive housing service firm is $C_2 = \eta H$, $\eta > 0$. Zero profit condition in turn gives $Z_t = \eta P_{2t}$, which is used to substitute out $\hat{Z}_{t+1}$ in (A3).

6.2 Renters

Without access to the housing investment market, a representative renter maximizes

$$\max E_0 \sum_{t=0}^{\infty} \beta^t \frac{1}{1-\frac{1}{\sigma}} u(C^r_{1t}, C^r_{2t})^{1-\frac{1}{\sigma}}, \quad 0 < \beta < 1, \quad \sigma > 0 \tag{A6}$$

subject to her budget constraint

$$Y^r_t + R_{t-1} B^r_{t-1} = P_{1t} C^r_{1t} + P_{2t} C^r_{2t} + B^r_t$$

where the superscript r denotes renters' choice variables.

The first order conditions for (C_1, C_2, B) are given by

$$\omega^{\frac{1}{\varepsilon}} [u^r_t]^{\frac{1}{\varepsilon}-\frac{1}{\sigma}} C^{r-\frac{1}{\varepsilon}}_{1t} = P_{1t} \mu^r_t \tag{A7}$$

$$(1-\omega)^{\frac{1}{\varepsilon}} [u^r_t]^{\frac{1}{\varepsilon}-\frac{1}{\sigma}} C^{r-\frac{1}{\varepsilon}}_{2t} = P_{2t} \mu^r_t \tag{A8}$$

$$\mu^r_t = \beta R_t E_t \mu^r_{t+1} \tag{A9}$$

After re-arranging log-linearized versions of (A7), (A8), and (A9), we get

$$\begin{aligned} & \Theta_1 \hat{P}_{1t} - \frac{1}{\sigma} \hat{C}^r{}_{1t} - \Theta_2 \hat{P}_{2t} \\ = \ & \hat{R}_t + \Theta_1 E_t \hat{P}_{1,t+1} - \frac{1}{\sigma} E_t \hat{C}^r{}_{1,t+1} - \Theta_2 E_t \hat{P}_{2,t+1}. \end{aligned} \tag{A10}$$

6.3 Aggregation

Supposing that the economy is inhabited by $\theta \in [0, 1]$ fraction of owner-occupiers and $1 - \theta$ fraction of renters, so that the aggregate consumption is determined by

$$C^A_t = \theta C_t + (1-\theta) C^r_t \quad . \tag{A11}$$

If log-linearized around a steady state, (A11) yields

$$\hat{C^A}_t = \Lambda\hat{C}_t + (1-\Lambda)\hat{C^r}_t \tag{A12}$$

where Λ is the steady state share $\theta C/C^A$ of owner-occupier's consumption.[11] We can further deduce that the same share Λ is applicable to the aggregation of C_1 or C_2 individually, in view of (A4) and the CES aggregator.

Linear combination of (A5) and (A10) gives

$$\begin{aligned} & \frac{1}{\sigma}\left[\hat{C^A}_{1t} - E_t\hat{C^A}_{1,t+1}\right] + \Theta_1\left[E_t\hat{P}_{1,t+1} - \hat{P}_{1t}\right] - \Theta_2\left[E_t\hat{P}_{2,t+1} - \hat{P}_{2t}\right] \\ = \ & \Lambda\left[\hat{Q}_t - \beta(1-\delta)E_t\hat{Q}_{t+1}\right] - (1-\Lambda)\hat{R}_t + \Lambda\left[1-\beta(1-\delta)\right]E_t\hat{P}_{2,t+1} \end{aligned} \tag{A13}$$

which we use to estimate σ and Λ.

[11] θ and Λ are different even in a steady state because the economy is populated by heterogenous households.

7 Appendix II : Derivation of Equation (10)

We consider another endowment economy in which the renters are, on top of the lack of access to the housing investment market, described as the so-called rule-of-thumb consumers: they do not smooth their consumption path in the face of fluctuations in their period-by-period endowment income. The behavior of home-owners in this economy are the same as in the Appendix II.

Each period a representative renter solves

$$\max \quad u(C^r_{1t}, C^r_{2t})$$

i.e., the static problem of maximizing his period utility subject to the constraint

$$Y^r_t = P_{1t}C^r_{1t} + P_{2t}C^r_{2t} \tag{A14}$$

that all his endowment income is consumed each period.

From the intratemporal optimization condition (A4) and the budget constraint (A14), we have

$$-\sigma \widehat{C^r}_{1t} = -\sigma \widehat{Y}^r_t + \sigma \widehat{P}_{1t}\left[\Theta_3 + (1-\Theta_3)\varepsilon\right] + \sigma \widehat{P}_{2t}(1-\Theta_3)(1-\varepsilon) \tag{A15}$$

where Θ_3 is the renters' share of non-housing consumption expenditure $P_1C^r_1/Y^r$ in steady state, which is equal to $\frac{\omega P_1^{1-\varepsilon}}{\omega P_1^{1-\varepsilon}+(1-\omega)P_2^{1-\varepsilon}}$[12].

We now combine the log-linearized versions of (A5) for owner-occupiers and (A15) for renters, with weight of Λ and $1-\Lambda$, respectively. Using

$$E_t\widehat{C}_{1t} = \frac{1}{\Lambda}E_t\widehat{C^A}_{1t} - \frac{1-\Lambda}{\Lambda}E_t\widehat{C^r}_{1t}$$

[12] We use the budget constraint (A14) and the CES aggregator to derive this reuslt.

we finally have the following equation for aggregate non-housing consumption:

$$
\begin{aligned}
\frac{1-\Lambda}{\sigma} E_t \widehat{Y^r}_{t+1} = \; & \Lambda \left[\hat{Q}_t - \beta(1-\delta) E_t \hat{Q}_{t+1} \right] \\
& + \Lambda \Theta_1 \left[\hat{P}_{1t} - E_t \hat{P}_{1,t+1} \right] - \Lambda \Theta_2 \left[\hat{P}_{2t} - E_t \hat{P}_{2,t+1} \right] \\
& + \Lambda \left[1 - \beta(1-\delta) \right] E_t \hat{P}_{2,t+1} - \frac{1}{\sigma} \left[\widehat{C^A}_{1t} - E_t \widehat{C^A}_{1t+1} \right] \\
& + \sigma(1-\Lambda) Y_t^r - \sigma(1-\Lambda) \left(\Theta_3 + (1-\Theta_3)\varepsilon \right) \hat{P}_{1t} \\
& - \sigma(1-\Lambda) (1-\Theta_3)(1-\varepsilon) \hat{P}_{2t} \\
& + \frac{1-\Lambda}{\sigma} \left(\Theta_3 + (1-\Theta_3)\varepsilon \right) E_t \hat{P}_{1,t+1} \\
& + \frac{1-\Lambda}{\sigma} (1-\Theta_3)(1-\varepsilon) E_t \hat{P}_{2,t+1}
\end{aligned} \tag{A16}
$$

which can be used an orthogonality condition for GMM estimation.