Wealth-to-Income Ratio, Housing Returns, and Systemic Risk

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Abstract

We show that the residuals of the trend relationship among asset wealth and human wealth predict housing returns. Using data for a set of industrialized countries, we assess the predictive ability of the wealth-to-income ratio for housing returns. In particular: (i) when housing asset are complements of financial assets, investors demand a higher housing risk premium if they are hit by a shock that generates a fall in the wealth-to-income ratio; (ii) when housing assets are substitutes of financial assets, investors demand a lower housing return if they face a fall in the wealth-to-income ratio. Finally, we show that the transmission of wealth shocks to housing markets is amplified in the outcome of episodes of systemic crises.

Keywords: Wealth; labour income; housing returns; systemic risk.
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1. Introduction

Differences in expected returns across assets are explained by differences in risk, and the risk premium is generally considered as reflecting the ability of an asset to insure against consumption fluctuations (Sharpe, 1964). Despite this belief, a measure such as the covariance of returns across portfolios and contemporaneous consumption growth did not prove to be sufficient to explain the differences in expected returns (Breeden et al., 1989). In fact, the literature on asset pricing has concluded that inefficiencies of financial markets[1] and the rational response of agents to time-varying investment opportunities[2] help justifying why expected excess returns appear to vary with the business cycle.

In addition, different macro-financially motivated variables that capture time-variation in expected returns have been developed. For instance: the consumption-wealth ratio (Lettau and Ludvigson, 2001); the long-run risk (Bansal and Yaron, 2004; Bansal et al., 2005); the labour income risk (Julliard, 2004); the housing collateral risk (Lustig and van Nieuwerburgh, 2005); the ultimate consumption risk (Parker and Julliard, 2005); and the composition risk (Yogo, 2006; Piazzesi et al., 2007); the ratio of excess consumption (i.e. consumption in excess of labour income) to observable assets (Whelan, 2008); and the wealth composition risk (Sousa, 2010a, 2010b). Similarly, for bonds, Silva et al. (2004) find that excess returns can be predicted by the Treasury yield spreads. Silva et al. (2003) also show that the inverse relative wealth and the dummy variable for the month of January are useful predictors of bond excess returns.

In contrast with the literature on the predictability of stock returns, only a few studies tried to explain the factors behind housing premia. Sousa (2010a) provides the first attempt to highlight this issue. In fact, the author shows that while financial wealth

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[1] See Fama (1998), Fama and French (1996), and Farmer and Lo (1999).

[2] See Sundaresan (1989), Constantinides (1990), Campbell and Cochrane (1999), Duffee (2005), and Santos and Veronesi (2006).

shocks are mainly transitory, fluctuations in housing wealth are very persistent. As a result, wealth composition might also be important because it has implications for the predictability of asset returns. In addition, De Veirman and Dunstan (2008) and Fisher et al. (2010) apply the approach developed by Lettau and Ludvigson (2001) to, respectively, New Zealand and Australia, and find the elasticity of consumption to permanent housing wealth changes is stronger than for permanent financial wealth variation. Sousa (2007) shows that housing can be used as a hedge against unfavourable wealth variation.

The current paper argues that the wealth and macroeconomic data can be combined to address the issue of predictability of housing returns for a set of industrialized countries. More specifically, we assess the forecasting power of the ratio of asset wealth to human wealth for expected future housing returns.

The rationale behind this linkage lies on the fact that a decrease in asset wealth reduces the value of collateral and increases household's exposure to idiosyncratic risk. Consequently, investors demand a higher stock risk premium when they face a fall in the ratio of wealth-to-income. However, in the case of housing returns, one needs to understand the way housing assets are perceived by agents. If they are seen as complements of financial assets, then investors behave in the same way as for stocks. However, if housing assets are substitutes of financial assets, then investors will require a lower housing risk premium when the ratio of wealth-to-income falls.

Using data for fifteen industrialized countries, we show that the ratio of aggregate wealth to income, *wy*, predicts housing returns, which helps understanding the importance of composition of asset wealth in the context of forecasting asset returns as Sousa (2010b) suggests.

The empirical findings suggest that the predictive power is especially important for horizons spanning from 4 to 8 quarters. In particular, the forecasting ability of *wy* for real housing returns is substantial, ranging between 1% (Australia, France and UK), 3% (Ireland), 7% (Germany and Netherlands), 15% (US), 16% (Sweden), 19% (Italy), 28% (Denmark), 38% (Belgium) and 41% (Spain) over the next 4 quarters. As for Canada, Finland and Japan, that proxy does not seem to capture well the time-variation in housing returns.

Moreover, the analysis suggests that one can cluster the set of countries into two groups. In the first group (which includes Denmark, Italy, UK and US), *wy* has an associated coefficient with negative sign in the forecasting regressions. Therefore, this corroborates the idea that housing and financial assets are complements in asset wealth. In the second group (which includes Australia, Belgium, France, Germany, Ireland, Netherlands, Spain and Sweden), the forecasting regressions show that *wy* has an associated coefficient that is positive. Consequently, agents in these countries understand housing assets as being substitutes for financial assets in their portfolios.

Finally, we ask whether the occurrence of systemic and non-systemic crises can amplify the transmission of wealth shocks to the housing market. We show that the predictive power of future housing returns is indeed improved when one takes into account the presence of crises' episodes, especially, the systemic ones.

The research presented in this work is related with the findings of Sousa (2010c) and Sousa (2010d). Using a set of sixteen industrialized countries, Sousa (2010c) shows that the residuals of the trend relationship among asset wealth and human wealth predict both stock returns and government bond yields. The author finds that when the wealth-to-income ratio falls, investors demand a higher risk premium for stocks. As for government bond returns: (i) when they are seen as a component of asset wealth, investors react in the same manner; (ii) if, however, investors perceive the increase in government bond returns as signalling a future rise in taxes or a deterioration of public finances, then investors interpret the fall in the wealth-to-income ratio as a fall in future bond premia. In the same context, Sousa (2010d) uses a panel of 31 emerging economies, and shows, from the consumer's budget constraint, that the residuals of the trend relationship among consumption, aggregate wealth, and labour income predict both stock and housing returns.

We improve upon the abovementioned papers in two major directions. First, we extend the analysis of the predictive ability of the wealth-to-income ratio to housing risk premium, while Sousa (2010c) focus on stock returns and government bond yields. Second, we develop a theoretical framework that highlights the role of wealth in the investors' utility function and also provide evidence for a set of industrialized countries. In this case, we depart from Sousa (2010d), who considers the representative agent's intertemporal budget constraint to derive an empirical proxy that captures time-variation in stock and housing returns, and builds on evidence for emerging market economies.

The paper is organized as follows. Section 2 describes the theoretical framework and the empirical approach. Section 3 presents the estimation results of the forecasting regressions for housing returns and the robustness analysis. Section 4 analyses the role of systemic risk in strengthening the linkages among the wealth-to-income ratio and housing returns. Finally, in Section 5, we conclude and discuss the implications of the findings.

2. Theoretical framework and empirical approach

2.1. Wealth-to-income ratio and (housing) risk premium

We follow Sousa (2010c) and assume that there is a continuum of agents who consume nondurable consumption, c_t, and wealth services (for instance, liquidity or collateral services), w_t, and are endowed with stochastic labor income, $y_t(i_t, a_t)$, where i_t represents the idiosyncratic event and a_t denotes the aggregate event.

The household maximizes utility, that is

$$U(c_t, w_t) = \sum_{s_t|s_0} \sum_{t=0}^{\infty} \beta^t p(s_t \mid s_0) u[c_t(s_t), w_t(s_t)], \tag{1}$$

where β is the time discount factor, s_t represents the state of the economy, $p(s_t \mid s_0)$ denotes the probability of state s_t given the initial state s_0, and preferences are specified by

$$u(c_t, w_t) = \left[c_t^{(\varepsilon-1)/\varepsilon} + \psi w_t^{(\varepsilon-1)/\varepsilon}\right]^{(1-\gamma)\varepsilon/(\varepsilon-1)} /(1-\gamma), \tag{2}$$

where ψ >0 captures the importance of wealth in the utility function, ε is the intratemporal elasticity of substitution between consumption and wealth services, and γ is the coefficient of risk aversion.

The solvency constraints are restrictions on the value of the household's consumption claim net of its labour income claim, that is:

$$\Lambda_{s_t}\left[c_t(s_t) + \rho_t(a_t) w_t(s_t)\right] \geq \Lambda_{s_t}\left[y_t(s_t)\right], \tag{3}$$

where $\Lambda_{s_t}[d_t(s_t)]$ represents the price of a claim to $d_t(s_t)$, and ρ_t is the rental price of wealth services.

The strength of the solvency constraints is determined by the ratio of asset wealth to human wealth (i.e., the wealth-to-income ratio), *wy*,

$$wy_t(a_t) = \Lambda_{z_t}\left[\rho w^a\right] / \Lambda_{z_t}\left[c^a\right]. \tag{4}$$

where w^a and c^a correspond, respectively, to aggregate wealth and aggregate consumption.

Equilibrium allocations and prices will depend on the consumption weight θ as follows: 1) if the household *does not switch* to a state with a binding constraint, it is $\theta_t'(\theta, s_t)$; and 2) if it *switches*, then the new weight is the cutoff level $\underline{\theta}_t(y_t, a_t)$.

In order to obtain aggregate consumption, one integrates over the new household weights, that is, $\zeta_t^a(a_t) = \int \theta'_t(\theta, s_t) d\Phi_t(\theta; a_t)$, where $\Phi_t(\bullet; a_t)$ represents the distribution over weights at the start of period *t*. The consumption share of an agent can then be represented as the ratio of his consumption weight to the aggregate consumption weight $c_t(\theta, s_t) = \theta'_t(\theta, s_t) \cdot c_t^a(a_t) / \zeta_t^a(a_t)$ and, similarly, for the wealth share of an agent $w_t(\theta, s_t) = \theta'_t(\theta, s_t) \cdot w_t^a(a_t) / \zeta_t^a(a_t)$, where $\zeta_t^a(a_t)$ defines a nondecreasing stochastic process.

As the ratio of wealth to income, *wy*, decreases, the cutoff levels for the consumption weights increase, $\underline{\theta}(y_t, a_t) / \zeta_t^a(a_t)$, and, if the consumer moves to a state where the constraint is binding, then the cutoff level for the consumption share equals the household's labour income share. As a result, when the ratio of wealth to income falls, the household's exposure to income shocks increases and a higher risk premium is demanded. Consequently, it should predict a rise in future stock returns.

Regarding housing returns, if housing assets are complements of stocks, then investors react in the same way. If, however, the increase of the exposure in risky assets is achieved via a fall of the share of wealth held in the form of housing (i.e., when stock and housing assets are substitutes), then they will demand a lower housing risk premium when they observe a fall in the wealth-to-income ratio. This behavior reflects the degree of separability between financial and housing assets: when they are separable, financial and housing assets will be substitutes, so agents can easily "smooth out" any transitory movement in their asset wealth arising from time variation in expected return; if, however, they are non-separable, financial and housing assets will be complements, and agents will not be able to "smooth out" exogenous shocks. Therefore, valuable information can be extracted by looking at the sign of the coefficients associated to *wy* in the forecasting regressions for housing returns.

2.2. Wealth, labour income, and housing returns

Log real per capita asset wealth, $log\ (w_t)$, and labour income, $log\ (y_t)$, are nonstationary. As a result, we estimate the following vector error correction model (VECM):

$$\begin{bmatrix} \Delta \log(w_t) \\ \Delta \log(y_t) \end{bmatrix} = \alpha\left[\log(w_t) + \varpi \log(y_t) + \vartheta t + \chi\right] + \sum_{k=1}^{K} D_k \begin{bmatrix} \Delta \log(w_{t-k}) \\ \Delta \log(y_{t-k}) \end{bmatrix} + \varepsilon_t, \quad (5)$$

where t denotes the time trend and χ is a constant. The K error correction terms allow one to eliminate the effect of regressor endogeneity on the distribution of the least-squares estimators of $\left[1, \varpi, \vartheta, \chi\right]$.

The components $log\ (w)$ and $log\ (y)$ are stochastically cointegrated and we impose the restriction that the cointegrating vector eliminates the deterministic trends, so that $\log(w_t) + \varpi \log(y_t) + \vartheta t + \chi$ is stationary. Then, the ratio of wealth to income, *wy*, is measured as the deviation from the cointegration relationship:

$$wy_t = \log(w_t) + \hat{\varpi} \log(y_t) + \hat{\vartheta} t + \hat{\chi}. \quad (6)$$

3. Results

3.1. Data

The data are quarterly, post-1960, and include fifteen countries (Australia, since 1970:1; Belgium, since 1980:2; Canada, since 1965:1; Denmark, since 1977:1; Finland, since 1979:1; France, since 1970:2; Germany, since 1965:1; Ireland, since 1975:4; Italy, since 1971:4; Japan, since 1965:1; the Netherlands, since 1975:1; Spain, since 1978:1; Sweden, since 1977:1; the UK, since 1961:2; and the US, since 1965:1). It, therefore, cover the last 30 to 50 years of data.

Labour income is approximated with compensation series of the NIESR Institute. In the case of the US, we follow Lettau and Ludvigson (2001). As for the UK, we follow Sousa (2010b).

Aggregate wealth data come from National Central Banks, the Eurostat, the Bank for International Settlements (BIS), the United Nation's Bulletin of Housing Statistics for Europe and North America.

Housing returns are computed using the housing price index and the price-rent ratio provided by the BIS and the Organization for Economic Co-Operation and Development (OECD). The dividend yield ratio is provided by Datastream.

Data for population are taken from OECD's Main Economic Indicators and interpolated from annual series.

Finally, all series are deflated, and expressed in logarithms of per capita terms and seasonally adjusted, with the obvious exceptions of housing returns.

3.2. The long-run relation

First, we use the Augmented Dickey and Fuller (1979) and the Phillips and Perron (1988) tests to determine the existence of unit roots in the series of aggregate wealth and labor income and conclude that they are first-order integrated, *I(1)*. Next, we analyze the existence of cointegration among the two series using the methodologies of Engle and Granger (1987) and Johansen and Juselius (1990), and find evidence that supports that hypothesis. Finally, we estimate the vector error-correction model (VECM) as expressed in (5).

Table 1 – Cointegration estimations. $wy_t = \log(w_t) + \hat{\varpi}\log(y_t) + \hat{\vartheta}t + \hat{\chi}$.

	$\hat{\varpi}$	Augmented Dickey and Fuller (1979) t-statistic	MacKinnon (1996) Critical values	
		Lags: Automatic based on Schwartz Information Criteria (SIC)	5%	10%
Australia	1.73*** (3.72)	-2.04	-2.88	-2.58
Belgium	1.06** (2.05)	-3.16	-2.88	-2.58
Canada	2.89*** (4.11)	-3.12	-2.88	-2.58
Denmark	-6.35* (1.87)	-2.88	-2.88	-2.58
Finland	2.17*** (12.53)	-2.73	-2.88	-2.58
France	1.04*** (3.05)	-2.68	-2.88	-2.58
Germany	0.63*** (2.76)	-3.78	-2.88	-2.58
Ireland	1.99*** (4.72)	-2.51	-2.88	-2.58
Italy	1.10*** (3.73)	-3.55	-2.88	-2.58
Japan	1.94*** (4.56)	-2.38	-2.88	-2.58
Netherlands	1.08** (1.92)	-3.43	-2.88	-2.58
Spain	4.60*** (4.71)	-2.64	-2.88	-2.58
Sweden	1.19* (1.56)	-2.17	-2.88	-2.58
UK	0.79* (1.36)	-2.31	-2.88	-2.58
US	0.53* (1.45)	-2.70	-2.88	-2.58

Notes: Newey and West (1987) corrected t-statistics appear in parenthesis. *, **, *** - statistically significant at the 10, 5, and 1% level, respectively.

Table 1 reports the estimates (ignoring coefficients for the constant and the trend) of the equilibrium relationship between aggregate wealth and labour income. First, it shows that the coefficient associated to income in the cointegrating vection is statistically significant for all countries, therefore, giving rise to the existence of an economically meaninful linkage between the two aggregates. Second, the point estimates for income are positive (with the exception of Denmark). This suggests that wealth and income tend to share a positive long-run path. Finally, the cointegration tests show that the residuals of the cointegrating relationship among aggregate wealth and income are stationary.

3.3. Forecasting housing returns

Section 2 shows that transitory deviations from the long-run relationship among wealth and income, wy_t, reflect agents' expectations about future housing returns.

We look at real housing returns (denoted by HR_t) for which quarterly data are available.

Note that long-horizon returns are calculated by summing the (continuously compounded) quarterly returns. This implies that the observations on long-horizon returns overlap which possibly biases the different test statistics towards rejecting the null hypothesis of no predictability more often than is correct (Nelson and Kim, 1993; Stambaugh, 1999; Valkanov, 2003; Ang and Bekaert, 2006). Nevertheless, one should emphasize that these works focus on the predictive ability of the dividend yield and the price-earnings ratio which are very persistent regressors. In contrast, we assess the forecasting power of the deviations from the equilibrium relationship between wealth and labor income, wy, which exhibit much less persistence. Thus, the abovementioned problems become less severe. Additionally, Lettau and Ludvigson (2001), Whelan (2008) and Sousa (2010b) find that the bias does not impact on the predictive ability of a wide range of variables in the forecasting regressions for stock returns. Finally, the adopted methodology is standard in the empirical finance literature (Lettau and Ludvigson, 2001; Julliard, 2004; Lustig and van Nieuwerburgh, 2005; Santos and

Veronesi, 2006; Yogo, 2006; Fernandez-Corugedo et al., 2007; Piazzesi et al., 2007; Sousa, 2010b).

Keeping these questions in mind, Table 2 summarizes the forecasting power of wy_t for different horizons. It reports estimates from OLS regressions of the H-period real housing return, $HR_{t+1} + \ldots + HR_{t+H}$, on the lag of wy_t. Therefore, we estimate the following model:

$$\sum_{h=1}^{H} HR_{t+h} = \alpha + \beta wy_{t-1} + \varepsilon_t \,. \qquad (7)$$

It shows that wy_t is statistically significant for a large number of countries, with the exceptions of Canada, Finland and Japan. Moreover, the trend deviations explain an important fraction of the variation in future real housing returns (as described by the adjusted R^2), in particular, at horizons spanning from 4 to 8 quarters. In fact, at the 4-quarter horizon, wy_t explains 1% (Australia, France and UK), 3% (Ireland), 7% (Germany and Netherlands), 15% (US), 16% (Sweden), 19% (Italy), 28% (Denmark), 38% (Belgium) and 41% (Spain) of the real housing return.

The results also suggest that the sign of the coefficient of wy_t is positive for Australia, Belgium, France, Germany, Ireland, Netherlands, Spain and Sweden, which therefore, indicates that agents demand a lower housing risk premium when they observe a fall in the wealth-to-income ratio. In this case, housing assets are seen as substitutes for financial assets. As for Denmark, Italy, UK and US, the sign of the coefficient of wy_t is negative and supports the idea that housing assets are complements of financial assets: when the ratio of asset wealth to human wealth falls, investors demand a higher risk premium for housing.

Table 2 – Forecasting real housing returns: estimated effect of *wy*.

	Forecast Horizon *H*						Forecast Horizon *H*				
	1	2	3	4	8		1	2	3	4	8
Australia	0.04**	0.06**	0.07*	0.06	0.02	Ireland	0.05	0.09	0.15	0.23*	0.20***
	(2.15)	(2.17)	(1.84)	(1.37)	(0.30)		(0.92)	(1.00)	(1.36)	(1.73)	(12.12)
	[0.03]	[0.03]	[0.02]	[0.01]	[0.00]		[0.01]	[0.01]	[0.02]	[0.03]	[0.03]
Belgium	0.06***	0.12***	0.18***	0.24***	0.51***	Italy	-0.28***	-0.51***	-0.68***	-0.81***	-1.44
	(5.43)	(7.33)	(9.21)	(9.68)	(11.57)		(-4.49)	(-4.66)	(-4.74)	(-4.93)	(-7.31)
	[0.20]	[0.31]	[0.35]	[0.38]	[0.47]		[0.24]	[0.23]	[0.21]	[0.19]	[0.30]
Canada	-0.00	-0.01	-0.00	0.00	0.05	Japan	0.02	0.00	0.02	0.04	-0.02
	(-0.38)	(-0.32)	(-0.20)	(0.06)	(0.94)		(0.30)	(0.03)	(0.32)	(0.70)	(-0.21)
	[0.00]	[0.00]	[0.00]	[0.00]	[0.01]		[0.00]	[0.00]	[0.00]	[0.01]	[0.00]
Denmark	-0.02***	-0.05***	-0.09***	-0.12***	-0.28***	Netherlands	0.05***	0.09***	0.12***	0.15***	0.27***
	(-2.49)	(-3.43)	(-3.89)	(-4.66)	(-6.73)		(4.41)	(4.48)	(4.41)	(4.23)	(4.53)
	[0.07]	[0.15]	[0.21]	[0.28]	[0.54]		[0.08]	[0.08]	[0.07]	[0.07]	[0.07]
Finland	0.06	0.10	0.13	0.16	0.31	Spain	0.08***	0.16***	0.23***	0.30***	0.42***
	(1.17)	(1.22)	(1.06)	(1.08)	(1.60)		(6.82)	(8.41)	(9.14)	(9.45)	(6.71)
	[0.02]	[0.02]	[0.02]	[0.02]	[0.03]		[0.28]	[0.38]	[0.41]	[0.41]	[0.26]
France	0.03**	0.05**	0.06*	0.05	-0.02	Sweden	0.06***	0.12***	0.17***	0.23***	0.40***
	(2.48)	(2.14)	(1.75)	(1.29)	(-0.29)		(2.94)	(4.70)	(5.40)	(6.69)	(5.97)
	[0.04]	[0.03]	[0.02]	[0.01]	[0.00]		[0.05]	[0.10]	[0.13]	[0.16]	[0.16]
Germany	0.04***	0.08***	0.12***	0.14***	0.16***	UK	0.02	0.01	-0.03	-0.09	-0.44***
	(3.06)	(4.15)	(4.78)	(5.32)	(3.82)		(0.60)	(0.17)	(-0.34)	(-0.82	(-3.15)
	[0.04]	[0.06]	[0.07]	[0.07]	[0.05]		[0.01]	[0.00]	[0.00]	[0.01]	[0.09]
						US	-0.03***	-0.07***	-0.11***	-0.14***	-0.27***
							(-4.68)	(-5.35)	(-5.96)	(-6.24)	(-6.31)
							[0.09]	[0.12]	[0.14]	[0.15]	[0.16]

Notes: Newey and West (1987) corrected t-statistics appear in parenthesis. Adjusted R-square is reported in square brackets. *, **, *** denote statistical significance at the 10, 5, and 1% level, respectively.

3.4. Nested forecast comparisons

Some recent studies (Bossaerts and Hillion, 1999; Goyal and Welch, 2003, 2004) expressed concerns about the apparent predictability of stock returns because, while a number of financial variables display significant in-sample forecasting power, they seem to have negligible out-of-sample predictive properties. In addition, the forecasting results presented so far could suffer from the "look-ahead" bias that arises from a long-term relationship estimated using the full sample (Brennan and Xia, 2005). In this context, some robust statistics such as the Clark and McCracken's (2001) encompassing test (ENC-NEW), the McCracken's (2006) equal forecast accuracy test (MSE-F) and the modified Diebold and Mariano (1995) encompassing test proposed by Harvey et al. (1998) could allow one to explore the out-of-sample performance of the forecasting model. Note, however, that the in-sample and the out-of-sample tests are equally reliable under the null of no predictability (Inoue and Killian, 2004). Moreover, the results from out-of-sample forecasts where the cointegrating vector is reestimated every period using only the data available at the time of the forecast could strongly understate the predictive power of the regressor (Lettau and Ludvigson, 2001). Therefore, it would make it difficult for *wy* to display forecasting power when the theory is true. Finally, Hjalmarsson (2006) shows that out-of-sample forecasting exercises are unlikely to generate evidence of predictability, even when the correct model is estimated and there is, in fact, predictability.

With these caveats in mind and as a final robustness check, we make nested forecast comparisons, in which we compare the mean-squared forecasting error from a series of one-quarter-ahead out-of-sample forecasts obtained from a prediction equation that includes *wy* as the sole forecasting variable, to a variety of forecasting equations that do not include it.

We consider two benchmark models: the *autoregressive benchmark* and the *constant expected returns benchmark*. In the *autoregressive benchmark*, we compare the mean-squared forecasting error from a regression that includes just the lagged housing return as a predictive variable to the mean-squared error from regressions that include, in addition, *wy*. In the *constant expected returns benchmark*, we compare the mean-squared forecasting error from a regression that includes a constant (as the only explanatory variable) to the mean-squared error from regressions that include, in addition, *wy*. As a result, the unrestricted model *nests* the benchmark model.

Table 3 summarizes the nested forecast comparisons for the equations of the real housing returns using *wy*. It shows that, in general, models that include *wy* generally have a lower mean-squared forecasting error. This is particularly important when the benchmark model is the *constant expected returns benchmark*, and, therefore, supports the existence of time-variation in expected housing returns.

Table 3 – One-quarter ahead forecasts of housing returns:
wy model vs. constant/AR models.

	Real housing returns	
	$MSE_{wy}/MSE_{constant}$	MSE_{wy}/MSE_{AR}
Australia	0.990	1.002
Belgium	0.898	0.970
Canada	1.003	1.003
Denmark	0.967	0.971
Finland	0.996	1.005
France	0.986	1.002
Germany	0.982	0.987
Ireland	1.001	1.005
Italy	0.876	0.931
Japan	1.003	0.999
Netherlands	0.965	1.004
Spain	0.856	0.984
Sweden	0.977	0.991
UK	1.001	0.991
US	0.959	0.986

Note: MSE – mean-squared forecasting error.

4. Does systemic risk matter?

Financial crises can be contagious and damaging, and prompt quick policy responses, as they typically lead economies into recessions and sharp current account imbalances. Among the many causes of financial crises, one can refer: (i) credit booms; (ii) currency and maturity mismatches; (iii) large capital inflows; and (iv) unsustainable macroeconomic policies (large current account deficits and rising public debt).

In order to deal with financial crises, governments have employed a broad range of policies, which reallocate wealth toward banks and debtors and away from taxpayers.

Honohan and Laeven (2005) and Laeven and Valencia (2008) identify financial crises episodes, and systemic crisis includes currency, debt and banking crises. A systemic currency crisis corresponds to a nominal depreciation of the currency of at least 30% and, simultaneously, at least a 10% increase in the rate of depreciation compared to the year before. A systemic debt crisis describes a situation where there are sovereign defaults to private lending and debt rescheduling programs. In a systemic banking crisis, there is a large number of defaults on corporate and financial sectors, non-performing loans increase sharply and, asset prices (equity and real estate prices) eventually depress, and real interest rates increase dramatically.

4.1. Systemic crises

In order to assess the importance of systemic crises, we estimate the following model:

$$\sum_{h=1}^{H} HR_{t+h} = \alpha + \beta wy_{t-1} + \mu wy_{t-1} * SystemicCrisis + \varepsilon_t , \qquad (8)$$

where *SystemicCrisis* is a dummy variable that takes the value of 1 in the presence of an episode of systemic crisis and 0 otherwise, and *H* refers to the number of ahead periods in the forecasting exercise. Given that the effects of systemic crises may not be immediate, we consider *H=1*, therefore, allowing for a time lag from the date of occurrence of the crisis and the emergence of its effects.

Table 4 – Forecasting real housing returns: impact of systemic crises.

	wy_{t-1}	wy_{t-1} * *SystemicCrisis*	Adj. R-square		wy_{t-1}	wy_{t-1} * *SystemicCrisis*	Adj. R-square
Australia	0.06***	-0.17***	[0.08]	Ireland	No episodes of systemic crisis		
	(-3.48)	(-3.74)					
Belgium	No episodes of systemic crisis			Italy	-0.32***	0.14*	[0.24]
					(-3.66)	(1.67)	
Canada	-0.00	-0.05	[0.00]	Japan	No episodes of systemic crisis		
	(-0.34)	(-1.12)					
Denmark	-0.02**	-0.00	[0.07]	Netherlands	No episodes of systemic crisis		
	(-2.37)	(-0.10)					
Finland	No episodes of systemic crisis			Spain	No episodes of systemic crisis		
France	0.03***	0.24***	[0.08]	Sweden	No episodes of systemic crisis		
	(2.53)	(3.52)					
Germany	0.03***	0.18***	[0.10]	UK	0.09**	-0.08	[0.06
	(2.80)	(3.13)			(2.07)	(-1.29)	
				US	-0.03***	0.07*	[0.08]
					(-4.58)	(1.62)	

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. *, **, *** - statistically significant at the 10, 5, and 1% level, respectively.

Table 4 reports the estimates from 1 quarter-ahead forecasting regressions. The results show that the point coefficient estimates of *wy* and their statistical significance do not change with respect to the previous findings. Moreover, the coefficient associated with the interaction between *wy* and the dummy variable for the systemic crisis is, in general, statistically significant.

4.2. Non-systemic crises

Finally, we analyse the impact of non-systemic systemic crises, and estimate the following model:

$$\sum_{h=1}^{H} HR_{t+h} = \alpha + \beta wy_{t-1} + \mu wy_{t-1} * NonSystemicCrisis + \varepsilon_t, \tag{9}$$

where *NonSystemicCrisis* is a dummy variable that takes the value of 1 in the presence of a non-systemic crisis and 0 otherwise, and *H* refers to the number of quarters-ahead of the forecasting exercise. Similarly to the case of systemic crisis, we allow for a lag in the transmission of the effects of non-systemic crises to housing markets and consider *H=1*.

Table 5 – Forecasting real housing returns: impact of non-systemic crises.

	wy_{t-1}	wy_{t-1} * *SystemicCrisis*	Adj. R-square		wy_{t-1}	wy_{t-1} * *SystemicCrisis*	Adj. R-square
Australia	No episodes of non-systemic crisis			Ireland	No episodes of non-systemic crisis		
Belgium	No episodes of non-systemic crisis			Italy	No episodes of non-systemic crisis		
Canada	No episodes of non-systemic crisis			Japan	-0.00	0.07	[0.01]
					(-0.03)	(0.64)	
Denmark	No episodes of non-systemic crisis			Netherlands	No episodes of non-systemic crisis		
Finland	-0.10	0.31***	[0.14]	Spain	No episodes of non-systemic crisis		
	(-1.40)	(3.04)					
France	No episodes of non-systemic crisis			Sweden	0.06***	0.08	[0.06]
					(2.51)	(0.60)	
Germany	No episodes of non-systemic crisis			UK	No episodes of non-systemic crisis		
				US	No episodes of non-systemic crisis		

Notes: Newey-West (1987) corrected t-statistics appear in parenthesis. *, **, *** - statistically significant at the 10, 5, and 1% level, respectively.

Table 5 summarizes the results from 1 quarter-ahead forecasting regressions. Again, the empirical evidence suggests that the point coefficient estimates of *wy* and their statistical significance remain unchanged. In what concerns the coefficient associated with the interaction between *wy* and the dummy variable for the non-systemic crisis, the results are somewhat weaker, especially, in comparison with the ones found for systemic crises. In fact, the interaction term is not statistically significant in most of the cases. However, its sign is typically positive, implying that, the occurrence of a non-systemic crisis leads investors to demand a higher risk premium for housing.

5. Conclusion

The current financial crisis has highlighted the strong connections between the financial system, the housing sector, and the banking sector not only in domestic terms, but also when considering inter-country dimensions. In fact, as Mallick and Mohsin (2007, 2010) note, monetary policy can be crucial, in particular, if it targets financial conditions (Castro, 2010; Sousa, 2010a).

This paper explores the predictive power of the trend deviations among asset wealth and human wealth (summarized by the variable *wy*) for future housing returns.

As in Sousa (2010c), the above-mentioned common trend summarizes agent's expectations of asset returns. In particular, when the wealth-to-income ratio falls (increases), forward-looking investors will demand a higher (lower) risk premium for stocks given that they will be exposed to larger (smaller) idiosyncratic shocks. Regarding housing returns, if housing assets are complements of financial assets, then investors behave in the same manner. If, however, housing assets are substitutes of financial assets, then investors will interpret the fall in the wealth-to-income ratio as predicting a decrease in future housing risk premium.

Using data for fifteen industrialized countries, we show that the predictive power of *wy* for real housing returns is particularly strong at horizons from 4 to 8 quarters. The analysis also suggests that one can consider two sets of countries: (i) those where housing assets are complements of financial assets (Denmark, Italy, UK and US); and (ii) those where agents see housing assets as substitutes for financial assets (Australia, Belgium, France, Germany, Ireland, Netherlands, Spain and Sweden).

Finally, we find that systemic crises amplify the effects of idiosyncratic shocks on housing markets. Consequently, the present work opens new avenues of investigation for understanding the dynamics of the relationship between wealth and housing market.

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