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WEALTH EFFECTS IN EMERGING ECONOMIES

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Abstract

In this paper I estimate the impact of changes in real and financial wealth – proxied by house and stock market prices – on private consumption for a panel of sixteen emerging economies in Asia and Central and Eastern Europe. Using recent econometric techniques for heterogeneous panels, i.e. the pooled mean group estimator, inference is drawn about the long- and short-run relationship between the variables of interest. Both real and financial wealth are found to affect household consumption positively in the long-run, with the elasticity of housing wealth being greater than that of stock market wealth. When the model is run separately for the two groups of countries, the long-run impact of an increase (decrease)in house prices is generally greater in Central and Eastern European economies than in Asian ones, which make the former more vulnerable to further adverse developments in the housing market.

JEL Classification: C23, D12, E20, E21, E32, E44, R21, R31.
Keywords: house prices, wealth effects, emerging markets, panel co-integration, pooled mean group estimator.

Contents

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1. Introduction [1]

Developments in housing markets have been attracting much attention in recent years, especially after the onset of the financial and real crisis erupted in September 2008 with the collapse of Lehman Brothers. The boom in house prices experienced by many countries in the years leading up to this adverse event, along with the subsequent busts, have been at the centre of both policymakers' and researchers' discussions, focused on the link between housing prices and the business cycle as well as the potentially adverse consequences on global financial stability.

Much of the analysis, nevertheless, has looked at the experience in advanced economies – United States and United Kingdom above all – while the characteristics of the housing markets, the forces driving prices as well as the links with overall macroeconomic conditions and the business cycle have not been yet systematically researched for emerging market economies (EMEs). Extending the existing literature to these countries is important not only because they are becoming a key engine of global growth, but also because in increasingly integrated capital markets the external financial spill-overs associated with adverse swings in their housing markets can be more serious. As a matter of fact, housing valuations in emerging regions have been rapidly catching up with those in the developed world, often fuelled by unprecedented expansions of credit to the private sector in the form of housing loans, and accompanied by sharp increases in leverage and exposures of households and financial intermediaries.[2] Real house prices have rapidly changed their course after the onset of the financial and real crisis of late 2008 sliding fast towards their long-run averages, with the largest declines recorded in those countries which have previously experienced the largest run-ups.[3]

Leaving aside the important theme of what potential factors might be at the roots of such dramatic shifts in housing valuations, i.e. good fundamentals vs. speculative forces,[4] the objective of the following analysis is rather to dig into one of the possible channels linking housing market developments and the business cycle in emerging economies, i.e. the relationship between house prices – and house price changes – and private consumption spending through the existence of a direct real (housing) wealth effect.

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[1] This research was finalized when I was visiting the Faculty of Economics of the University of Cambridge. I am greatly indebted to Francesco Bripi, Luisa Corrado, Antonio De Socio, Valeria Rolli, Vanessa Smith, Giorgio Trebeschi, Teng Teng Xu and two anonymous referees for useful comments on earlier versions of this paper; any errors and omissions remain my own responsibility. The usual disclaimers apply.

[2] Led by some countries in Central and Eastern Europe – like Latvia, Lithuania, Poland, Russia and Estonia – average real house prices increased by almost 50 percent in real terms from 2005 – chosen as a base year due to the relatively mild movements in housing valuations – to the most recent peak in early 2008. Asian economies, on the contrary, have all shown below-the-average movements in house prices during this time span, with Singapore and Hong Kong recording the highest increases in real terms (42.4 and 26.6 percent, respectively) and Korea, Malaysia and Thailand experiencing the lowest (11, 2.7 and 0.6 percent, respectively). Nevertheless, all the emerging economies considered in this research have shared a similar upward trend with advanced countries, corroborating the thesis of an increasing coincidence or real house prices movements internationally (Girouard *et al*., 2006).

[3] Housing valuations, in fact, have come down substantially in Latvia, Estonia, Lithuania and Bulgaria where, at their latest available trough, they were 60, 58, 51 and 41 percent lower in real terms than the previous peak, respectively. The smallest declines have been recorded in Thailand, Malaysia and Korea, where real house prices were, at their respective trough, 7.3, 4.4 and 4.2 percent lower than the previous peak.

[4] The interest reader is referred to Ciarlone (2010) for an analysis of the determinants of house price cycles in emerging economies.

As the provided evidence shows for a sample of 16 emerging economies, households consumption, disposable income and two measures of real and financial wealth – proxied, respectively, by house and stock prices – are found to be difference-stationary and co-integrated; by means of recent econometric techniques designed for heterogeneous panels (Pesaran *et al.*, 1999), a reduced-form consumption function is estimated, which offers evidence of the existence of both a real (housing) and a financial (stock market) wealth effect, with the former being larger than the latter. Moreover, comparing the coefficients estimated in this paper with those presented in the extant literature on advanced economies would seem to suggest that the housing wealth effect turns out to be larger in the emerging countries with respect to the developed ones, while the opposite holds true for the stock market wealth effect.

Developments in housing markets are therefore found to be able to affect the business cycle in emerging economies through their effects on private consumption spending. This leaves open the relevant question as to which policy levers are best suited to deal with boom-bust cycles in housing markets: monetary, fiscal and macro-prudential policies are all expected to play a potential role in this respect. Unfortunately, the policy agenda is still in the phase of exploration (Crowe *et al*., 2011), and much further research would be therefore needed.

My analysis contributes to the empirical literature in several ways. First of all, it updates previous research focusing on emerging economies (Peltonen *et al.*, 2009) by taking into account the most recent period of financial turmoil (up to the end of 2009), when collapses in house prices endured prolonged distress in several countries. Second, it applies up-to-date econometric techniques, specifically engineered for heterogeneous co-integrated panels, in order to analyze the relationship between consumption and the two wealth components, housing vs. stock market: by doing so, I would be able to reach more rigorous results with respect to those presented in other, rather scant, pieces of empirical literature on the same topic.

The paper is structured as follows: Section 2 deals with the theoretical underpinnings behind the relationship between consumption, real (housing) and financial (stock market) components of wealth, and briefly reviews the extant empirical literature on housing wealth effects; Section 3 delves into the main features of the econometric procedure, the pooled mean group estimator proposed by Pesaran *et al.* (1999); Section 4 and 5 host, respectively, the description of the data and the results of the preferred econometric specification, clearly showing the existence of a housing wealth effect which clearly outweighs the stock market one. Section 6, at last, summarizes the main conclusions of the paper.

2. House prices, wealth effects and consumption

2.1 Housing and the macro-economy

Housing market developments are supposed to affect a country's macroeconomic conditions through different channels.

First of all, house prices – and house price changes – may impinge negatively on financial stability (Hilbers *et al.*, 2008). Booms in housing valuations have been shown to have explanatory power in early warning systems for financial crises, increasing the probability that the eventual turning point in prices might be followed by rising tensions in the banking sector (Bunda and Ca' Zorzi, 2010).

Apart from these more extreme events, house prices – and house price changes – are supposed to influence a country's aggregate demand and growth prospects through their effects on investment and consumption spending.

As regards supply-side effects, they are indeed more readily apparent, since house price movements may directly influence the 'residential' component of a country's gross fixed capital formation. The construction sector, a significant contributor to value added, takes property prices as a signal and adjusts production accordingly: big corrections in housing construction may therefore have a non negligible impact on employment and growth. Residential investments appear to lead the business cycle – at least in many advanced economies, but also in China – and a softening of housing construction may be an important factor leading to a cyclical downturn.

As regards the demand-side effects, they are debated from a theoretical point of view. A growing piece of empirical literature has been suggesting the existence – at least for advanced economies – of a possibly relevant impact of changes in house prices – and housing wealth – on private consumption and saving decisions, representing the main channel through which the business cycle is thereby affected. As a matter of fact, for the available sample of 16 emerging economies, changes in residential property prices – used, in what follows, as a proxy for changes in housing wealth – and changes in households consumption display a predictable pattern over the last decade or so, with a correlation of almost 0.6 (**Chart 1**);[5] moreover, co-movements between the two variables seem to have strengthened recently, especially after the onset of the real and financial crisis in late 2008. For the majority of developing countries, residential property represents households largest asset:[6] therefore, for a shock of similar magnitude, the wealth effect of changes in house prices would be larger than for other financial asset prices. An additional more indirect factor works through the credit channel, at least for emerging economies with more sophisticated housing finance systems: although changes in housing prices may be considered just a redistribution of wealth, and hence would not be expected to have much impact on net wealth in aggregate, they can nevertheless affect individual consumption by relaxing collateral constraints (Buiter, 2008).[7]

[5] A similar result has been found by Catte *et al.* (2004), for a sample of advanced economies.

[6] This is testified by the high home-ownership rates which, according to the most recent available data (ranging between 2004-2007), are estimated to be around 59 percent in emerging Asia, 62 percent in Latin America and 82 percent for Central and Eastern Europe (Peltonen *et al.*, 2009; European Mortgage Federation, 2007; national statistical offices). Moreover, these values are comparable with those for developed economies: for instance, home-ownership rate hovers around 71 percent in the UK, 65 percent in the US and 56 percent in France (Crowe *et al.*, 2011).

[7] Consistent with this point, changes in house prices have shown to have a medium-run liquidity effect on both US and UK consumption, once the impact of credit market developments and deregulation – which raises access to housing collateral – is carefully taken into account (Muellbauer, 2008).

2.1 Housing and private consumption

A preliminary issue to tackle is to ascertain the channels through which changes in house prices could lead to changes in households wealth. From a theoretical point of view, the answer is not so straightforward. The conceptual framework which tends to be used to analyze consumption dynamics is the permanent income hypothesis (PIH) developed by Friedman (1957) and the steady-state form of the life-cycle hypothesis (LCH) developed by Modigliani and Brumberg (1954) and Modigliani (1975), according to which consumption spending is determined by expected lifetime resources. Within this framework, it is important to take into account the fact that households both own housing assets and consume housing services deriving from them.[8] Following a house price increase, therefore, home-owners may feel being wealthier through both a realized wealth effect – since it is possible for them to take out equity in the form of refinancing or selling the house – or an unrealized wealth effect – since, even if households do not sell their house, they can be expected to spend more today due to a higher discounted value of wealth. An increase in house prices, nonetheless, also implies a rise in the value of housing services, thereby generating a budget constraint effect on both home-owners and home-renters which works in the opposite direction with respect to both the realized and unrealized wealth effects. In sum, the capital gains to the home-owners may end-up being partly, or fully, offset by the higher discounted values of future imputed rents, i.e. by the higher opportunity cost of owning vs. renting a house, therefore not implying any increase in net wealth and consumption.

Moreover, unlike a rise in equity prices which can reflect an increase in the economy's expected potential growth and future income prospects, higher house prices may simply mirror scarcity owning to higher demand, with no change in either the quantity or the quality of the services housing can provide to the overall economy. Nevertheless, even if aggregate wealth is unchanged, house price increases usually affect the relative position of specific groups of people, i.e. current home-owners vs. would-be buyers, and these wealth transfers may have macroeconomic effects if these categories' propensity to spend differs.[9]

Finally, another feature of the housing market to be taken into account is its illiquidity. Compared with other financial assets, it is relatively costly and time-consuming to convert increases in housing wealth into money which can be directly spent. Consequently, the response of consumers to house price shocks can be qualitatively different from the reaction to financial asset price shocks. In particular, consumer spending would respond to a house price shock only after the accumulated price movement become so large that it exceeds the transaction costs associated with adjusting the housing stock.

Notwithstanding these theoretical uncertainties and ambiguities, the assessment of housing wealth effects has conveyed a substantial empirical literature. Much of the work on this topic delves into the

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[8] The price of housing services reflects the amount of money that tenants spend on the provision of shelter (rents) and that owner occupants would have spent had they been renting (imputed rental value).

[9] Would-be buyers are typically people acquiring a house early in life, when their income is relatively low; since the purchase of a house typically requires a large sum of money, which rises with its value, many young households may actually save more when real estate prices increase.

experience of advanced economies, mainly the US and UK, and makes use of flows-of-funds aggregate data (which are available for longer periods, higher frequency and great timeliness than household-level data). Results are usually summarized in terms of the marginal propensity to consume out of housing wealth, which indicates how much consumption changes (in absolute terms) for a one dollar/euro change in housing wealth; other works, instead, estimate the consumption response to housing wealth shocks in terms of elasticities.[10] In both cases estimation results seem to point to the existence of a positive and significant effect of housing wealth on consumption (see ECB, 2009, for an up-to-date review of all the empirical works that have reached this conclusion); this effect, moreover, would seem to be larger for economies characterized by more developed financial markets and housing finance systems (Ludwig and Slok, 2002; Catte *et al.* 2004; Buiter, 2008; Muellbauer, 2008).

There is much less consensus, instead, on how the housing wealth effect compares with the financial wealth effect, often proxied by changes in stock market valuations. For the sample of countries at hand, stock prices show a certain degree of co-movement with households consumption, with a correlation of almost 0.5 (**Chart 2**). The theoretical reasons for such differences are not clear-cut. On the one hand, shocks to the financial wealth are typically more transitory than shocks to housing wealth, which would suggest a smaller propensity to consume out of financial wealth; on the other hand, the transaction costs of realizing capital gains from favourable asset price movements are relatively much higher for housing assets, and this fact could lead to the opposite conclusion. These ambiguities are reflected by the results of the empirical literature: while some studies find that the housing wealth effect is substantially stronger than the financial wealth effect (e.g. Case *et al.*, 2001; Case *et al.*, 2011; Carrol *et al.*, 2006; Pichette and Tremblay, 2003; Slacalek, 2006), others report the opposite (Bassanetti and Zollino, 2008; Calomiris *et al.*, 2009; Dreger and Reimers, 2009; De Bonis and Silvestrini, 2012; Ludwig and Slok, 2002 and Ludwig and Slok, 2004).

While all the above papers have analyzed the relationship between housing wealth and private consumption in advanced countries, very few studies exist on emerging economies, as severe problems of data availability hamper a complete and effective empirical assessment. Funke (2004) was the first – to my knowledge – to show some evidence of a small, but statistically significant, wealth effect in a sample of 16 emerging economies, but its analysis was confined to stock market wealth effects. Building upon this result, Peltonen *et al.* (2009) estimated the magnitude of a stock market, a housing and a money wealth effect on consumption in a panel of 14 emerging economies; according to their results, all these effects are statistically significant and relatively large, with the one from money wealth being the largest. The latter paper is similar to mine in terms of the broad analysis, while the main difference is the preferred econometric methodology. Their model, in fact, relies upon techniques developed in the context of traditional dynamic panels with a relatively large N (the number of cross-sectional observations) and small T (the number of time-series observations) dimension, i.e. the difference-GMM estimator originally

[10] Of course the two measures are strictly interrelated, since the elasticity of consumption to housing wealth equals the marginal propensity to consume multiplied by the ratio of housing wealth to consumer expenditure.

pioneered by Arellano and Bond (1991) and further improved by Blundell and Bond (1998). While these procedures essentially require pooling individual countries and allowing only the intercepts to differ across them, one of the central findings of the more recent literature, which focuses on dynamic panels in which both N and T are large and possibly of the same order of magnitude (as it is the case in Peltonen *et al.*'s paper), is that the assumption of homogeneity of slope parameters is appropriate only under quite strong assumptions.[11] Moreover, traditional procedures for estimation of pooled models – such as fixed-effects, instrumental variables and GMM estimators – can produce inconsistent and biased estimates of the average values of the parameters, unless the slope coefficients are in fact identical as pointed out in Pesaran and Smith (1995), Pesaran *et al*. (1999) and Im *et al*. (2003).

In what follows, I will present new improved estimates of the real and financial wealth effects on households consumption for a sample of 16 emerging economies by means of a recent econometric procedure that allows for weaker homogeneity assumptions in the context of possibly non-stationary panels, the pooled mean group (PMG) estimator pioneered by Pesaran *et al*. (1999).

3. The econometric model

A simple model of an aggregate consumption function with household (labour) income and wealth as the only determinants is theoretically founded on Friedmans' PIH and Modigliani's LCH.[12] In most empirical analysis of the wealth effect on consumption, a common trend among the three variables is assumed and tested for: Galì (1990), for instance, provides a theoretical foundation for a common trend approach between these three macroeconomic aggregates. A potential econometric problem in estimating a consumption function pertains to the correlation between consumption and the components of wealth: when seeking to identify the effect of an increase in wealth on consumption, in fact, the estimated conditional correlations may end up reflecting, to a certain extent, the effect of an increase in consumption on wealth. This 'reverse causality' is traditionally referred to as endogeneity bias: failing to address this problem could skew the statistical inference, and lead to inconsistent estimates of how much an increase in wealth influences consumption.

In what follows it is assumed – and tested if – a co-integrating relationship exists between consumption, income, and the two measures of financial and real wealth; next, an error-correction specification of a consumption function is estimated for the sample of 16 EMEs.

The literature on inference in dynamic and co-integrated panels has evolved rapidly over the past decade, proposing a number of methods which are designed not only to deal with the endogeneity bias correction, but also to accommodate for nuisance parameters and serial correlation in data. Some procedures are based on vector error-correction representation (Mark and Sul, 2003; Breitung, 2005), while

[11] In particular, it is required that the group-specific parameters are distributed independently of the regressors, and that the regressors are strictly exogenous.

[12] Altissimo *et al.* (2005) and Davidson *et al.* (1978) demonstrated these points for the PIH and LCH, respectively.

others are based on single-equation approaches, like the fully-modified OLS estimator (Pedroni, 2000), the PMG estimator (Pesaran *et al.*, 1999) and the dynamic OLS (DOLS) estimator (Kao and Chiang, 2000).

In this paper, I will implement two single-equation parametric estimators, and compare the results obtained from each of them. The first approach is given by the PMG specification proposed by Pesaran *et al*. (1999), which allows for very flexible assumptions on the panel framework under consideration: specifically, the pooled estimation of consumption elasticities with respect to income, financial and real wealth is carried out by imposing long-run restrictions, while the short-term parameters, the speed of adjustment and the innovations variances are left unrestricted across countries. Moreover the PMG estimator, building upon the auto-regressive distributed lag (ARDL) modelling approach to co-integration analysis developed by Pesaran and Shin (1997), can simultaneously correct for serial correlation in the residuals and the problem of endogenous regressors through an appropriate choice of the lag structure of both the dependent and the explanatory variables. The second approach – which will be used here mainly for the purpose of robustness check – is the panel DOLS estimator suggested by Kao and Chiang (2000), where a sufficient number of leads and lags of the first difference of the right-hand side variables is included in the estimation procedure to get rid of the effects of regressor endogeneity on the distribution of the OLS estimator.

Turning to the specification of the empirical model, an identical reduced-form of the long-run consumption function linking consumption, income and the two measures of wealth is assumed for all countries:

$$(1)\quad c_{t,i} = \alpha_{0,i} + \beta_{1,i} y^{d}_{t,i} + \beta_{2,i} w^{hw}_{t,i} + \beta_{3,i} w^{sw}_{t,i} + \varepsilon_{t,i} \qquad i=1,2,\ldots,N \quad t=1,2,\ldots,T$$

where c is households consumption, y^d is households income, w^{hw} and w^{sw} respectively refers to the real and financial components of households wealth and ε_{ti} is the error term capturing the effects of unexpected shocks to consumption; the subscripts t and i denote time and country, respectively.

An important issue to tackle is to give the previous representation a dynamic structure. There are various reasons for such a representation, such as the presence of habit persistence, adjustment costs, liquidity constraints and so on: all these aspects prevent immediate adjustment of consumption to a change in its fundamental determinants, and should be therefore taken into account in the empirical modeling. Influential literature, for instance, typically estimates ARDL models of consumption over income and wealth, introducing lag mechanisms to model the response of the former to changes in the latter variables.

Equation (1) could be therefore properly generalized by introducing deterministic terms, an auto-regressive lag polynomial for the dependent variable and complicated distributed lag schemes for the explanatory variables. For the ease of presentation it will be assumed here – but relaxed afterwards – that just the first lag of each variable is an important determinant of the consumption function in each country;[13] the ensuing ARDL (1,1,1,1) specification of equation (1), therefore, becomes:

$$(2)\ c_{t,i} = \alpha_{0,i} + \gamma_i c_{t-1,i} + \beta_{10,i} y^d_{t,i} + \beta_{11,i} y^d_{t-1,i} + \beta_{20,i} w^{hw}_{t,i} + \beta_{21,i} w^{hw}_{t-1,i} + \beta_{30,i} w^{sw}_{t,i} + \beta_{31,i} w^{sw}_{t-1,i} + \varepsilon_{t,i}$$

The error term is assumed to be independently distributed across t and i, but the variances may be heterogeneous across countries;[14] moreover, it is also supposed that the error term is independent of all the other variables in equation (2), an assumption that can be invalidated if other important determinants of households consumption are omitted from the empirical specification.

It is also necessary, at this point, to take into account another problem that may arise from equation (2): Pesaran and Shin (1997), in fact, have shown that the ARDL approach is no longer applicable whenever the variables in (2) are integrated of order 1. This problem could be easily overcome by simply re-parameterizing equation (2) in order to take into account the possible long-run relationship among the variables:

$$(3)\ \Delta c_{t,i} = \alpha_{0,i} + \varphi_i(c_{t-1,i} - \alpha_{1,i} y^d_{t-1,i} - \alpha_{2,i} w^{hw}_{t-1,i} - \alpha_{3,i} w^{sw}_{t-1,i}) + \beta_{10,i}\Delta y^d_{t,i} + \beta_{20,i}\Delta w^{hw}_{t,i} + \beta_{30,i}\Delta w^{sw}_{t,i} + \varepsilon_{t,i}$$

where

$$(4)\quad \varphi_i = -(1-\gamma_i); \qquad \alpha_{1,i} = (\beta_{10,i} + \beta_{11,i})/(1-\gamma_i);$$
$$\alpha_{2,i} = (\beta_{20,i} + \beta_{21,i})/(1-\gamma_i); \qquad \alpha_{3,i} = (\beta_{30,i} + \beta_{31,i})/(1-\gamma_i)$$

Equation (3) stands for the error-correction re-parameterization of the ARDL (1,1,1,1) model. Since, according to the Engle and Granger's representation theorem (1987) there is a clear link between co-integration and error-correction mechanism, equation (3) represents the starting point to carry out the estimation of the long-run relationship between consumption, income, real and financial wealth.

[13] Higher order lags can be easily accommodated within the same framework.

[14] I acknowledge the fact that the assumption of cross-sectional independence of the error term is indeed rather strong and restrictive, as macro time-series may, in some instances, exhibit significant degree of cross-correlation among the countries in the panel. This may arise because of the presence of common shocks and unobserved components that ultimately become part of the error term, spatial dependence and so on: one reason for this is given by the ever-increasing degree of economic and financial integration of countries and financial entities observed during the last few decades, which has gone along with a dramatic rise in the inter-dependencies between cross-sectional units. Cross-section dependence has been representing a rapidly growing field of study in panel data analysis, primarily aimed at investigating solutions to the adverse consequences on existing research instruments. The impact of cross-sectional dependence on dynamic panel estimators may indeed be quite severe: for instance, one of the most striking effects is that the pooled OLS estimator would provide little gain in precision compared with single equation OLS when cross-sectional dependence occurs in the data but is ignored in the panel regression, or that commonly used panel unit root tests would no longer be asymptotically similar (Phillips and Sul, 2003); or that standard IV and GMM estimators would suffer large rises in bias and RMSE if the amount of cross-section dependence is not small (Sarafidis and Robertson, 2006); or tests that suppose independence across cross-sectional units would suffer from size distortions (Moon and Perron, 2003). There is not enough room here to do justice to the ever-growing body of research on this topic; the interested reader is referred to Baltagi (2005) for a useful introduction.

In a panel setting as this, country-specific effects can be controlled for, for example, by using a dynamic fixed-effect (DFE) specification, which can be estimated using standard least-squares dummy variable (LSDV) or GMM estimators. However, DFE specifications typically impose homogeneity of all slope coefficients, allowing only the intercept to vary across countries. Denoting with k the number of long-run parameters, DFE imposes $(N-1)*(2k+2)$ restrictions (including the short-run dynamics) on the restricted model in equation (3), i.e. k long-run coefficients, k short-run coefficients plus the convergence coefficient and the common variance. The validity of the DFE, in particular, depends critically on the assumption of common convergence parameters which implies, on its turn, a certain degree of homogeneity among the units of the sample that is difficult to recognize in this particular case: Pesaran and Smith (1995) show that, under slope heterogeneity, both LSDV and GMM DFE estimators of the speed of convergence are usually affected by a downward heterogeneity bias. Consequently, DFE estimators do not appear suitable for implementation in this setting.

An alternative strategy would be to adopt a mean group (MG) estimator, which consists of estimating separate regressions for each country and calculating averages of country-specific coefficients (Pesaran and Smith, 1995). Albeit consistent, the estimator is likely to be inefficient in small country samples, where any country outlier could severely influence the averages of country coefficients.

Following Pesaran *et al.* (1999), I take an intermediate path between imposing ex ante homogeneity on all slope coefficients (as in the DFE case) and imposing no restrictions at all (as in the MG estimator case) by using the PMG estimator approach: it essentially requires allowing the intercepts, the speed of adjustment, the short-run coefficients and error variances to differ across countries, while imposing long-run parameters to be identical across groups, i.e. the 'long-run homogeneity' assumption termed by the authors. In other words, I am imposing $(N-1)*k$ restrictions on the unrestricted model shown in equation (3) of the type $\alpha_{.i} = \alpha$ for every i. The maximum likelihood estimator of the restricted equation is usually called the PMG estimator, and it has been shown to be asymptotically normal for the case of both stationary and non-stationary regressors.

4. Data description and preliminary tests

The data set consists of an unbalanced panel of 16 main emerging economies, 10 from Central and Eastern Europe (Bulgaria, the Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Russia, Slovakia and Slovenia) and 7 from Asia (China, Hong Kong, Korea, Malaysia, Singapore and Thailand); data are quarterly and span (when available) the period 1995Q1-2011Q2.

Given the broad coverage of this study, I encountered some data limitations. First of all, real and financial wealth derived from the flow of funds accounts – which might provide a more precise measurement of households total wealth (De Bonis and Silvestrini, 2012) – is not available on a broad basis for the sample of EMEs; wealth components, therefore, have been proxied by stock market and house price indexes.[15] Second, my focus here is on total consumption, and there is no distinction between the non-durable and durable components: again, a problem of data availability and homogeneity prevented me from following the conventional approach, i.e. to use non-durable consumption.[16] Third, gross national disposable income (GNDI) is used in the preferred specification instead of labour income, as would be suggested by the traditional permanent income hypothesis: data availability was again the constraint, but I will nevertheless turn to available labour income data as a robustness check.

Data on house prices have been collected from the BIS Data Bank, national central banks and statistical offices websites, Datastream and Bloomberg. Notwithstanding their limitations, a big effort has been made in making all the series as comparable as possible: annual data have been linearly interpolated using a quadratic matching average procedure; property prices in national currency have been transformed in nominal indexes, and all the nominal indexes have been rescaled to the same base year (2005); finally, all the re-based nominal indexes have been deflated by the country's CPI in order to express them in real terms.

Data on stock market indexes expressed in local currency have been collected from Datastream, therefore calculating their relative quarterly averages, rescaling them to the same base year of house prices, and expressing them in real terms.

Data on private consumption and GNDI have been collected from the IMF-International Financial Statistics database; in those instances where GNDI was not available, I supplanted it with the series of gross domestic product (GDP). Private consumption, GNDI and GDP are expressed in real, per capita terms on a seasonally-adjusted basis.

As a final remark, logs have been taken of all the variables, such that the estimated coefficients should be interpreted as elasticities of consumption with respect to changes of the right-hand side regressors: unfortunately, since comparable data for real and financial wealth in domestic currency are not available, it has not been possible to turn elasticities into marginal propensities to consume out of the two components of wealth.

The long-run equilibrium relationship between consumption, disposable income and real and financial wealth cannot be consistently estimated if all the single variables have unit roots, unless the variables in the long-run relationship are co-integrated. Therefore, one has to ascertain the statistical properties of the series, and test whether a co-integrating equilibrium relationship between households consumption, disposable income and the two proxies of real and financial wealth indeed exists.

[15] A high correlation has been traditionally found between stock market prices and wealth measures, as documented in Lettau and Ludvigson (2001), while ECB (2009) attests that developments of housing wealth in the euro area over the past 25 years generally borne a close resemblance to the evolution of euro area residential property prices.

[16] According to some pieces of literature (Romer, 1990; Mehra, 2001), total consumption is the parameter of interest when studying movements in stock market prices: crashes, in fact, are more likely to lead to a postponement of durable consumption, while the reduction of non-durable consumption might be of minor importance.

As a first step, therefore, I carried out a battery of first-generation panel unit root tests, i.e. the Breitung (2000), the Levin-Lin-Chu (2002), the Im-Pesaran-Shin (2003) and the Hadri (2000) tests. The results, hosted in **Table 1**, are clear-cut and imply not rejection of the presence of a unit root for almost all the series of interest, assuming a constant in the test regression. Only those obtained for real stock market prices may seem, at first sight, quite odd especially from the point of view of the efficient market hypothesis, which is in fact associated with the idea of a random walk.[17] Two aspects are, in my opinion, responsible for this outcome. On the one hand, running these tests requires a strongly balanced panel, where the time dimension is therefore constrained by that of the country with the shortest available series.[18] On the other hand, the financial crisis erupted in September 2008 brought about a sharp drop in stock valuations for all the emerging economies in the sample, ranging from 40 to 86 percent in real terms from peak to trough. Such a huge price fall, if studied within a panel characterized by a relatively small T dimension, may well end up being interpreted as a sort of 'mean-reverting process' by the unit root tests at hand. In order to shed more light on this aspect, I restricted the sample to those emerging economies with the longest available series (from 1994Q1 onwards),[19] and re-run all the previous tests: with the only exception of the Im-Pesaran-Shin test, all the others seem now to point more clearly to the random walk behaviour implied by the efficient market hypothesis.[20]

Once reassured by the results of panel unit root tests, I turned to the issue of testing for the existence of a co-integrating relationship among households consumption, disposable income, financial and real wealth. Like panel unit root tests, panel co-integration tests are motivated by the search for more powerful procedures than those obtained by applying individual time-series co-integration tests. While these latter tests, in fact, are known to have low power especially for short *T*, panel co-integration tests have the notable feature of being implementable with much shorter time spans of data, improving upon the small sample limitations of conventional non-stationary methods (Pedroni, 2000).

Like in standard time series, in a panel setting there are different ways of testing the null hypothesis of no co-integration. Typically, in fact, these tests are grouped in two large families, the residual-based and the likelihood-based tests: the former are constructed on the basis of the Engle and Granger’s (1987) test in time-series framework, and use residuals of the panel static regression to construct the test statistics and to tabulate the relative distributions; the latter, instead, represent generalization in the panel setting of the work pioneered by Johansen (1991, 1996) for vector auto-regressive models.

[17] According to this fundamental hypothesis, in efficient markets the flow of information is unimpeded and immediately reflected in stock prices; as a consequence, tomorrow's price changes will reflect only tomorrow's news, and will be independent of the price changes today. But news is, by definition, unpredictable and, thus, resulting price changes must be unpredictable and random as well. As a matter of fact, from an empirical point of view there is no consensus as to whether stock prices are mean-reverting or random walk processes: at best, the results are mixed.

[18] In this case Bulgaria, whose stock prices are available only from 2000Q4 onwards.

[19] The following countries have been excluded because of their relatively shorter series: Bulgaria, Estonia, Latvia and Lithuania.

[20] The results are not reported here for the sake of brevity, but are available from the author upon request.

Within the residual-based family, Kao (1999) studied a family of Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) tests for the null of no co-integration,[21] and derived their limiting distributions when applied to spurious regressions in a panel setting under the rather strong hypothesis of homogeneous co-integrating vectors between the sample units, i.e. not allowing for coefficient heterogeneity. He showed that, after appropriate normalizations, these test statistics converge, by sequential limit theory, to random variables with normal distributions. Skipping all the technical details – the interested reader is referred to the reference above and to Barbieri (2006) for an excellent review of panel co-integration tests – it suffices here to simplify by saying that Kao's approach requires first to estimate the presumed long-run relationship by pooled OLS, obtain the residuals, and finally implement a (normalized) pooled Dickey-Fuller (or Augmented Dickey-Fuller) regression on these residuals; the tests statistics, moreover, may contain nuisance parameters to account for possible weak exogeneity in the regressors and serial correlation in the residuals. **Table 2** displays the results for three test statistics of the Kao's family – the DF_t, the DFT_t* and the *ADF* – which clearly suggest rejection of the null of no co-integration.[22] I can be fairly confident of these results since it has been shown (Gutierrez, 2003) that, in homogeneous panels, Kao tests outperform, in terms of power, other residual-based panel co-integration tests when the time dimension of the panel is relatively small, as in this setting.

Since the hypothesis of coefficient homogeneity may seem rather strong, I submitted the previous results to a robustness check by employing the four new panel co-integration tests recently developed by Westerlund (2007). These tests are engineered to verify again the null hypothesis of no co-integration, but are based on structural – rather than residual – dynamics; as such, they are aimed at inferring whether the error-correction term in a conditional error-correction model is equal to zero.[23] They have been shown to be remarkably flexible, being able to accommodate for country-specific short-run dynamics, including serially-correlated error terms, non-strictly exogenous regressors, country-specific intercepts and trend terms, and country-specific slope parameters. Two statistics are designed to test the alternative hypothesis that the panel is co-integrated as a whole (P_α and P_τ), while the other two test the alternative that there is at least one country for which the variables are co-integrated (G_α and G_τ); asymptotic results reveal that they have limiting normal distribution, and that they are consistent. **Table 3** contains the value of the four test statistics, along with their *p*-values: all the results do point to the same direction, providing further support to the existence of a co-integrating relationship among the variables of interest.[24]

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[21] More precisely, Kao tests assume as null and alternative hypothesis that either all the relationship are not co-integrated or all the relationship are co-integrated.

[22] The main difference between the *DF* and the *DF** families relies on the hypothesis regarding the exogeneity of the regressors and the serial correlation in the residuals. More precisely, the DF* test statistics contain nuisance parameters to account for these two problems.

[23] If the null hypothesis of no error correction is rejected, then the null hypothesis of no co-integration is also rejected.

[24] As a matter of fact, in order to test for co-integration I also resorted to the methodology developed by Pedroni (2004), who proposed seven panel test statistics which allow for heterogeneity of the long-run covariance matrix for each unit i and of the slope parameters across all units i. Three out of these seven statistics – i.e. the Panel v, the Panel ADF and the Group ADF – again lead to reject the null hypothesis of no co-integration providing a further, although a bit less robust, check to the outcomes contained in the main text. Results are not reported here for the sake of brevity, but are available from the author upon request.

5. Estimation results

In summary, the available evidence seems to suggest that the variables of interest are non-stationary and co-integrated. Therefore, estimation of equation (3) with variables expressed in log levels provides reliable inferences about the long- and short-run influences of stock market and house prices and disposable income on consumption.

Since OLS estimators are super-consistent in case of co-integrated variables, the first estimates presented are static fixed-effects for the whole sample of countries, which are reported in **Table 4** along with the usual battery of unit root tests on the residuals of the estimated equation. The point estimates of the elasticities of consumption to changes in disposable income and stock and house prices are positive and statistically significant using robust standard errors, with the latter only at the ten percentage significant level.

Though giving a first signal about the existence of both a stock market and a housing wealth effect, traditional static panel techniques such as the former are nevertheless based on rather strong homogeneity assumptions among countries, by imposing a single slope coefficient in the pooled estimation. The assumptions underlying static panel techniques appear to be too stringent in the case under study: potential country heterogeneity, in fact, could be modelled in a much richer way than using simple fixed (or random) effects model. That's why I turned to the preferred econometric procedure, the PMG estimator.

As a first necessary step, for each country of the sample the lag order of the ARDL model has been chosen by applying the Schwarz information criterion, the results of which are reported in **Table 5**. Although there is no clear evidence of a most common representation, some general tendencies could nevertheless be found out by simply looking at the frequencies by which the lagged variables appear in each country's ARDL representation: for instance, stock prices and house prices appear more often than not in no lagged form, while the most frequent lag for both real consumption and real income is the first one. According to this evidence, the preferred specification for the whole sample of countries will be an ARDL (1,1,0,0) which in this framework will read as:

$$(5)\quad c_{t,i} = \alpha_{0,i} + \gamma_i c_{t-1,i} + \beta_{10,i} y^d_{t,i} + \beta_{11,i} y^d_{t-1,i} + \beta_{20,i} w^{hw}_{t,i} + \beta_{30,i} w^{sw}_{t,i} + \varepsilon_{t,i}$$

that is, real consumption is lagged once as real disposable income, while only contemporaneous terms of real house and real stock prices are included. Equation (5) can be re-parameterized as follows:

$$(6)\quad \Delta c_{t,i} = \alpha_{0i} + \varphi_i(c_{t-1,i} - \alpha_{1,i} y^d_{t,i} - \alpha_{2,i} w^{hw}_{t,i} - \alpha_{3,i} w^{sw}_{t,i}) - \beta_{10,i} \Delta y^d_{t,i} + \varepsilon_{t,i}$$

which represents the preferred specification to be estimated by means of the PMG estimator.

Table 6 contains the regression results, along with the respective *p*-values, for both the combined sample as well as for the two separate groups of EMEs, Asian and Central and Eastern European countries.[25] In the two regional regressions, I have repeated the exercise of choosing the ARDL representation according to the lag frequencies: for the Central and Eastern European economies, the preferred specification is still an ARDL (1,1,0,0), while it is an ARDL (1,2,0,0) for the Asian ones. Both the estimated house price and the stock market price elasticities turned out to be positive and significant: interestingly, the estimated size of the coefficient measuring the stock market wealth effect (0.02) is much smaller (almost a fifth) than the size of coefficient estimate on house price elasticities (0.10). Splitting the sample into the two regional groupings reveals that the estimated coefficients for Central and Eastern European countries are roughly similar to the group-wide results as regards the stock market wealth effect (0.02), while the elasticity to changes in housing prices is significantly higher (0.14 vs. 0.10); for Asian economies, instead, estimation results point to a lower impact on consumption from a shock to house market valuations with respect to the group-wide effect (0.06 vs. 0.10), while the opposite holds true for a shock to stock prices (0.04 vs. 0.02).[26]

The hypothesis of homogeneity of long-run coefficients implied by the PMG procedure cannot be assumed *a priori*, but needs to be tested. In order to do that, I followed the approach suggested by Pesaran *et al*. (1999), which essentially requires comparing the two sets of coefficient estimates obtained by means of the PMG and the MG procedures: as anticipated in Section 3, the MG estimator does not pose any kind of constraint on the coefficients, since it requires running *N* separate regressions and then averaging the parameters of interest; the PMG estimator, on the other side, pools the data and constrains some parameters (i.e. the long-run coefficients) to be the same across groups. When the long-run homogeneity restrictions are indeed true, the PMG procedure would yield consistent and more efficient estimates with respect to the MG approach (Pesaran et al., 1999); if the true model is heterogeneous, instead, the PMG estimates would be inconsistent. On the other side, the MG estimates would be consistent in either case. Based on this reasoning, it is possible to test the long-run homogeneity hypothesis by means of a Hausman test (Hausman, 1978): under the null, the difference in the estimated coefficients between the unrestricted (consistent) MG and the restricted (efficient) PMG specification is not significantly different from zero. The lower part of Table 6 hosts the values of the test statistic, which is distributed as a chi-squared with as many degrees of freedom as the parameters to estimate (also reported): the corresponding *p*-values seem to suggest that the null hypothesis of homogeneity restriction on long-term parameters cannot be rejected, and that moving from the MG to the PMG estimator is indeed appropriate for the data at hand.[27]

[25] These results have been obtained by means of the a-do Stata files developed by Blackburn III and Frank (2007).

[26] Results for the single countries in the sample have not been reported here for the sake of brevity, but are available from the authors upon request. It suffices here to say that the general conclusion of positive stock market and housing wealth effects do hold even in the case of individual countries, of course with differing degrees of elasticities.

[27] The coefficient estimates obtained by means of the MG procedure are not reported here for the sake of brevity, but are available from the author upon request.

The lower part of **Table 6** also reports the adjustment coefficients φ of the short-term dynamics for the entire sample and for the two country groupings:[28] they all have the correct negative sign – which implies that an error-correction mechanism is in place – and are always statistically significant. The estimation results for the regional groupings seem to suggest, moreover, that once the economy is hit by a shock, a slower adjustment process of consumption occurs in Central and Eastern European countries with respect to Asian ones, since the φ coefficient turns out to be lower for the former than for the latter group.

The general result, pointing to a positive responsiveness of consumption to changes in both wealth components – with the sensitivity to changes in housing wealth being larger than that to changes in stock market wealth – has been subjected to a number of robustness checks. In the first one, I modified equation (3) with the introduction of a deterministic time trend; in the second type of check, I estimated the consumption function with only one component of wealth each time; in the third test, I modified the definition of households income by substituting data relative to the GNDI from the IFS database with the series relative to the wages and salaries obtained from national sources (even if this substitution would introduce another source of heterogeneity, since the definition of the group of workers for which this statistics is produced, for instance, greatly differs from one country to the other); lastly, in the fourth test, I re-estimated the model by running the panel dynamic OLS procedure (Kao and Chiang, 2000), which essentially requires to add a sufficient number of leads and lags of the independent variables to take into proper account the endogeneity of the regressors and the serial correlation in the error term. The results of these four robustness checks – which are hosted in Table 7 under the headings Check 1, Check 2, Check 3 and Check 4 – seem all to confirm the general conclusion of the existence of both a stock market and a housing wealth effect on consumption, with the latter being larger than the former.[29] Finally, the Hausman tests run for each of these specifications seem again to suggest that imposing homogeneity restrictions on the long-run coefficients cannot be rejected, and that moving from the MG to the PMG estimator is again appropriate.

According to the different estimation procedures and model specifications, the elasticity of consumption to changes in housing wealth ranges between 0.06 and 0.20, with a mid-point of 0.13, while the elasticity of consumption to changes in stock market wealth ranges between 0.01 and 0.07, with a mid-point of 0.04. Overall these findings – i.e. the existence of both a housing and a stock market wealth effect, with the former being larger than the latter – are coherent with other pieces of empirical literature, even if any comparison should be taken *cum grano salis* because of the enormous differences in terms of frequency of data, sample of countries, definitions of wealth, time periods, estimation procedures, and so on. In the only other paper on emerging economies, with a sample which is partly different from mine, Peltonen et al. (2009) estimated a long-run house price elasticity in a range from 0.04 to 0.15 for Asia, and equal to

[28] The adjustment coefficients reported in the table are calculated as an un-weighted average of each country's ones, which are left to be heterogeneous by the PMG procedure. The single country's estimated adjustment coefficients are available from the author upon request.

[29] Note that, once one component of wealth is taken away from the estimation procedure, the coefficient of the disposable income approaches one, as implied by the life-cycle hypothesis.

almost 0.32 for Latin American countries, compared to a stock market elasticity equal to almost 0.09 independently from the empirical specification. As regards developed economies, the bunch of studies which define the wealth effects in terms of elasticities – and, for this reason, are more directly comparable to the present one – all support the conclusion of a positive co-movement between private consumption and housing wealth, while the evidence on the relationship between housing and stock market wealth effect is more mixed.[30] In extreme synthesis, this empirical evidence locates the elasticity of consumption to changes in housing wealth in an interval from 0.02 to 0.19, with a mid-point of 0.10, and the elasticity of consumption to changes in stock market wealth in an interval from 0.01 to 0.23, with a mid-point of 0.12. As a consequence, my estimates would seem to suggest that the housing wealth effect turns out to be larger in emerging economies with respect to advanced ones, while the opposite holds true for the stock market wealth effect.

6. Conclusions and policy recommendations

In this work, I have estimated the impact of changes in real and financial wealth on consumption for a panel of 16 emerging economies. Since flow of fund accounts data were not largely available, they have been supplanted with stock market and house price indexes, which have been used as proxies for real and financial wealth. Households consumption, income and the two measures of wealth have been found to be difference-stationary and co-integrated; by means of recent econometric techniques for heterogeneous panels, i.e. the pooled mean group estimator, inferences were drawn about the long- and short-run relationship between the variables of interest.

The main result of the analysis shows that both real and financial wealth positively affect households consumption in the long-run, with the elasticity of housing wealth being larger than that of financial wealth. According to the different estimation procedures, the long-term elasticity of consumption to changes in house prices ranges from 0.06 to 0.20, with a mid-point of 0.13; the elasticity of consumption to changes in stock market prices ranges from 0.01 to 0.07, with a mid-point of 0.04. If compared with the evidence available for advanced economies, these estimates would seem to suggest that the housing wealth effect turns out to be larger in emerging economies with respect to developed ones, while the opposite holds true for the stock market wealth effect. There is also a significant short-run adjustment from income, stock prices and house prices on consumption, i.e. consumption adjusts to its long-run relationship with lags, albeit at a quite fast rate: according to the different estimation procedures and model specifications, in fact, the average half-time of the adjustment process is approximately 3 quarters among the countries considered.

[30] While Case *et al.* (2001), Case *et al.* (2011) and Chen (2006) point to a housing wealth effect being larger than a stock market wealth effect, Bassanetti and Zollino (2008), Bertaut (2002), Boone *et al.* (2001), Dreger and Reimers (2009), Ludvig and Slok (2002), Ludvig and Slok (2004) all reach the opposite conclusion.

When the model is run separately for the two groups of economies, Asia vs. Central and Eastern Europe, it turns out that the elasticity of consumption to changes in housing prices is larger for the latter, while the adjustment process after the economy has been hit by a shock is slower. This implies that the countries in Central and Eastern Europe are more vulnerable to adverse developments in the housing sector, should the contraction in real house prices continue at the recently observed rates.

The reported evidence of a possibly relevant impact of changes in house prices – and housing wealth – on private consumption in emerging economies opens the way to the relevant question as to which policy levers are best suited to deal with housing market boom-bust cycles. In principle, monetary, fiscal and macro-prudential policies are all expected to play a dramatic role in this respect (Crowe *et al.*, 2011). An increase in the policy rates, for instance, would make borrowing more expensive, therefore reducing the demand for loans by households and the leverage in the financial sector. Nevertheless, monetary policy may turn to be a too blunt instrument: it may entail a substantial cost in terms of output gap and unemployment rate when the boom is limited to the real estate market; moreover, higher interest rates may encourage speculative capital inflows from abroad thereby further relaxing domestic liquidity conditions, a problem particularly relevant for emerging economies. As regards fiscal policy, an increase in transaction or property taxes, as well as reforms aimed at reducing the mortgage interest tax deductibility, can dampen volatile house price dynamics and the build-up of vulnerabilities associated with debt-financed ownership. Unfortunately, the available empirical evidence on the efficacy of such tools is, again, inconclusive; moreover, the scope for their use is complicated by the inherent difficulties related to the acceptance by the public opinion. Macro-prudential tools, such as higher capital requirements, dynamic provisioning and limits to loan-to-value and debt-to-income ratios, appear at first sight as best suited to achieve the objective of protecting the economy against adverse developments in real estate markets: notwithstanding some shortcomings, like the relative ease to be circumvented and the difficulty of implementation from a political economy standpoint, they have an undoubted ability to attack the problem at its source along with their added benefit of increasing the resilience of the banking system. But part of this reasoning is just speculation: the policy agenda is still in the phase of explorations, and much research would be therefore needed on this particular front.

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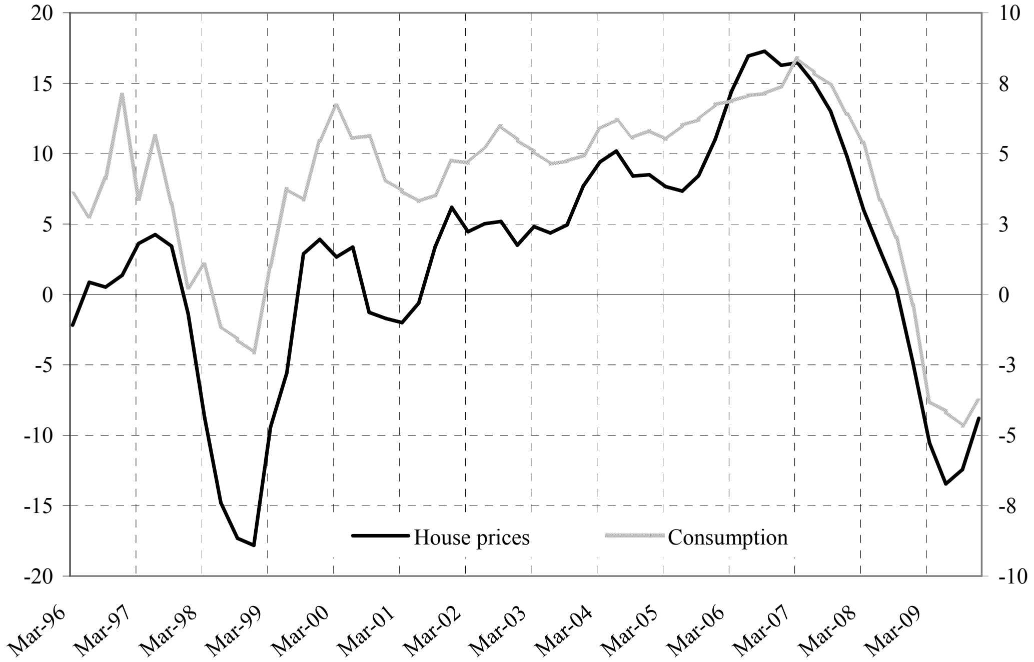
Tables and charts

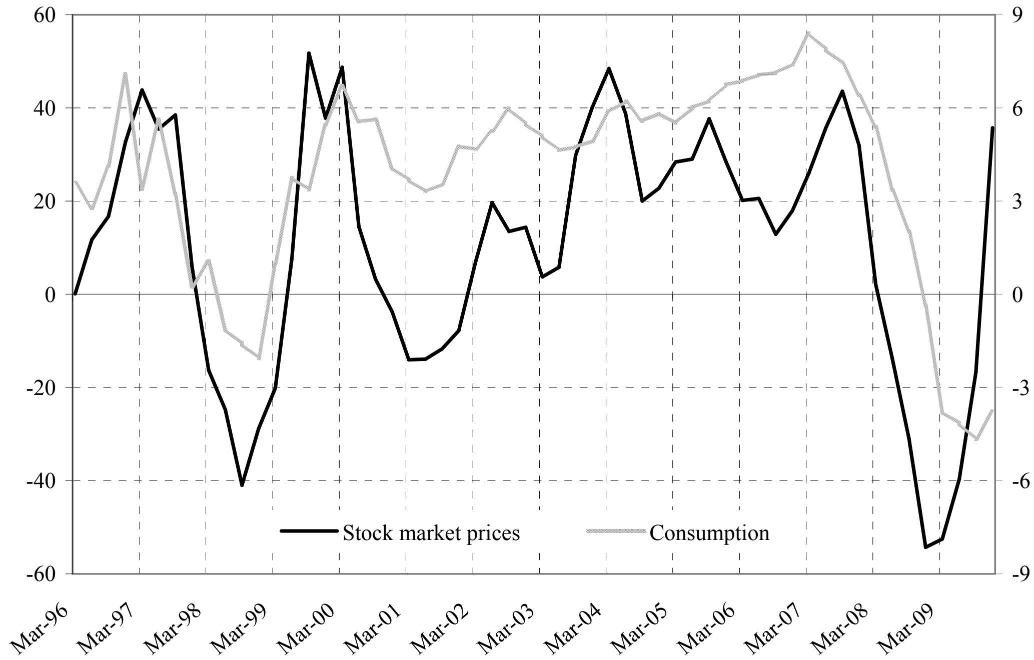
Chart 1. Real house prices and real private consumption in emerging economies
(quarterly data; average of annual percentage changes)



Note: the lines represent the simple average of the quarterly growth rates of real house price and real private consumption series for the 16 emerging economies in the sample.

Source: author's calculation on data from national central banks and statistical offices, BIS Data Bank, Thomson Reuters Datastream, Bloomberg, IMF-International Financial Statistics.

Chart 2. Real stock market prices and real private consumption in emerging economies
(quarterly data; average of annual percentage changes)



Note: the lines represent the simple average of the quarterly growth rates of real stock price and real private consumption series for the 16 emerging economies in the sample.

Source: author's calculation on data from national central banks and statistical offices, BIS Data Bank, Thomson Reuters Datastream, Bloomberg, IMF-International Financial Statistics.

Table 1. Panel unit root tests

Variable	Breitung	Levin-Lin-Chu	Im-Pesaran-Shin	ADF Fisher $\chi 2$	ADF Choi Z-statistic	Hadri
real consumption	0.08	3.68	5.21	16.96	4.90	19.86
	(0.53)	*(1.00)*	*(1.00)*	*(0.99)*	*(1.00)*	*(0.00)*
real disposable income	5.54	-2.84	1.76	36.07	1.88	19.42
	(1.00)	*(0.00)*	*(0.96)*	*(0.37)*	*(0.97)*	*(0.00)*
real house prices	1.27	-2.42	0.15	36.03	0.10	10.77
	(0.90)	*(0.01)*	*(0.56)*	*(0.37)*	*(0.54)*	*(0.00)*
real stock market prices	-2.20	-2.41	-2.35	45.74	-2.38	8.75
	(0.01)	*(0.01)*	*(0.01)*	*(0.09)*	*(0.01)*	*(0.00)*

Note: all p-values (in parenthesis) are reported such that Ho is rejected if p-value<0.05; models include an intercept.

Table 2. Panel co-integration: Kao tests (1999)

	DF type
DFt	-2.05
	(0.02)
DFt*	-2.25
	(0.01)
lags	ADF type
1	-5.29
	(0.00)

Note: Null hypothesis (Ho) is the estimated equation is not co-integrated. All p-values (in parenthesis) are reported such that Ho is rejected if p-value<0.05

Table 3. Panel co-integration: Westerlund tests (2007)

	t-statistic
$G\tau$	-2.54
	(0.00)
$G\alpha$	-16.43
	(0.00)
$P\tau$	-10.42
	(0.00)
$P\alpha$	-12.30
	(0.00)

Note: Null hypothesis (Ho) is the estimated equation is not co-integrated. All p-values (in parenthesis) are reported such that Ho is rejected if p-value<0.05

Table 4. Estimating consumption wealth effects: preliminary evidence based on static fixed effects *(dependent variable: log of real consumption)*

Variable	Coefficient
constant	0.51
	(0.00)
real disposable income	0.80
	(0.00)
real house prices	0.08
	(0.00)
real stock market prices	0.01
	(0.07)
Total panel (unbalanced) observations	799
Cross-sections included	17
Adjusted R^2	0.95
Breitung	-2.59
	(0.00)
Levin, Lin, Chu	-6.12
	(0.00)
ADF-Fisher χ^2	124.01
	(0.00)
ADF-Choi Z-statistic	-7.72
	(0.00)

Note: all p-values (in parenthesis) are reported such that Ho is rejected if p-value<0.05; unit root tests do not include an intercept nor a time trend.

Table 5. Auto-regressive distributed lag specification

	Real consumption	Real disposable income	Real stock prices	Real house prices
Bulgaria	1	1	0	0
China	2	2	0	0
Czech Rep.	2	1	0	0
Estonia	2	0	0	0
Hong Kong	1	1	1	2
Hungary	1	0	0	0
Korea	1	1	0	1
Latvia	2	0	1	0

	Real consumption	Real disposable income	Real stock prices	Real house prices
Lithuania	1	1	1	0
Malaysia	1	0	2	0
Poland	1	1	0	0
Russia	1	1	0	0
Singapore	2	2	0	0
Slovakia	1	1	0	0
Slovenia	1	0	0	0
Thailand	1	2	0	1

Note: orders of lags in the ARDL model selected by the Schwarz Information Criterion.

Table 6. Estimating consumption wealth effects: the pooled mean group estimator *(dependent variable: log of real consumption)*

Variable	**All countries**	**Asia**	**Central and Eastern Europe**
constant	0.21	0.29	0.15
	(0.00)	*(0.00)*	*(0.00)*
real disposable income	0.76	0.72	0.71
	(0.00)	*(0.00)*	*(0.00)*
real house prices	0.10	0.06	0.14
	(0.00)	*(0.03)*	*(0.00)*
real stock market prices	0.02	0.04	0.02
	(0.02)	*(0.00)*	*(0.03)*
Adjustment coefficient φ	-0.37	-0.35	-0.23
	(0.00)	*(0.00)*	*(0.00)*
Total panel (unbalanced) observations	799	406	470
Cross-sections included	16	6	10
Hausman test for long-run homogeneity			
degrees of freedom	3	3	3
value	6.64	4.96	2.58
	(0.08)	*(0.17)*	*(0.46)*

Note: all p-values (in parenthesis) are reported such that Ho is rejected if p-value<0.05; for Asia, ARDL(1,2,0,0); for Central and Eastern Europe, ARDL(1,1,0,0)

Table 7. Estimating consumption wealth effects: robustness checks *(dependent variable: log of real consumption)*

Variable	Check 1: deterministic trend	Check 2: only stock prices	only house prices	Check 3: wages and salaries	Check 4: panel DOLS
constant	0.32	0.12	0.19	0.25	0.54
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*
real disposable income	0.66	0.91	0.79	0.57	0.79
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*
real house prices	0.12		0.10	0.20	0.08
	(0.00)		*(0.00)*	*(0.00)*	*(0.00)*
real stock market prices	0.03	0.02		0.07	0.02
	(0.00)	*(0.02)*		*(0.00)*	*(0.00)*
trend	0.00				
	(0.03)				
Adjustment coefficient φ	-0.37	-0.35	-0.36	-0.32	-0.10
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*
Total panel (unbalanced) observations	799	951	799	663	799
Cross-sections included	16	16	16	16	16
Hausman test for long-run homogeneity					
degrees of freedom	4	2	2	3	
value	9.10	5.01	2.08	4.16	
	(0.06)	*(0.08)*	*(0.35)*	*(0.24)*	

Note: all p-values (in parenthesis) are reported such that Ho is rejected if p-value<0.05.

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