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The Impact of Changes in Stock Prices and House Prices on Consumption in OECD Countries

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The Impact of Changes in Stock Prices and House Prices on Consumption in OECD Countries

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Abstract

This paper quantifies the different impact of stock and house prices on consumption using data for 16 OECD countries. The analysis finds that the long-run impact of an increase in stock prices and house prices is in general higher in countries with a market-based financial system. The sensitivity of consumption to changes in stock wealth is about twice as large as the sensitivity to changes in housing wealth. Splitting the sample into the 1980s and 1990s shows that both countries with a market-based financial system and countries with a bank-based financial system moved toward a higher degree of responsiveness of consumption to changes in stock prices and house prices.

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I. Introduction

This paper examines the role of stock prices and house prices as determinants of private consumption. The impact of stock market wealth on consumption has received increased attention among both academic researchers and policymakers, partly because of the dramatic increase and subsequent fall in stock prices experienced in the 1990s (Figure 1).[2] Measured in percent of GDP the Anglo-Saxon countries have experienced the strongest gains (Figure 2) and have consequently also been the focus of most studies. For many continental European countries the increase in stock market wealth has also been quite substantial. In Japan, however, asset prices have fallen through the 1990s, reflecting the protracted downturn for the Japanese economy.

While movements in financial wealth have been dominated by movements in stock market wealth, housing wealth is the single most important component of non-financial wealth in households' portfolios (Deutsche Bank (2001)). Due to pronounced increases in housing wealth (Figure 3) and deregulation of mortgage markets, the impact of housing wealth on consumption in OECD countries has therefore also received increased attention among researchers and policymakers.[3] As Greenspan (2001) has recently suggested, the marginal propensity to consume out of housing wealth might dollar for dollar be higher than the marginal propensity to consume out of stock market wealth while the overall impact of the latter must expected to be higher.

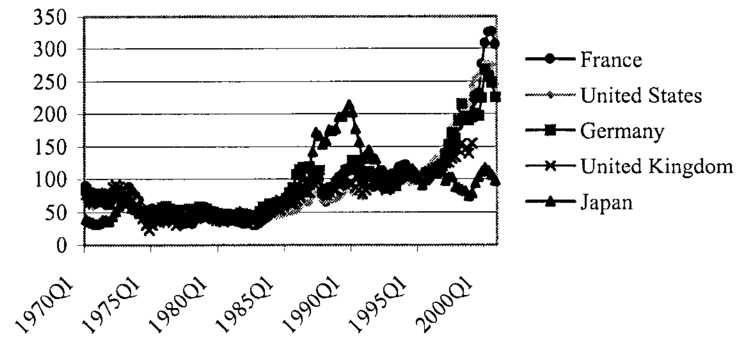
The views on the role between asset prices and real economic activity differ widely in the literature. At one end of the spectrum it has been argued that the observed correlation between asset prices and consumption expenditures are due to the role of asset prices as "leading indicators" (see e.g. Morck et al. (1990); Poterba and Samwick (1995)). According to this view, asset prices reflect future output growth and are therefore correlated with consumption expenditures. At the other end of the spectrum is the view that the observed correlation is due to real wealth effects. The distinction between these two perspectives is of high importance regarding policy conclusions from the observed correlations. It will therefore be further discussed in the concluding remarks in Section V.

This paper contributes to the existing literature by taking a broader perspective in investigating the relative importance of two wealth components – housing and stock market wealth – using quarterly data for a panel of 16 OECD countries. Countries are grouped into bank-based and market-based economies and differences between the relative importances of the two wealth components in these sub-groups are analyzed. A new panel data technique for

[2] See IMF (2000) and (2001a), Greenspan (2001), Edison and Sløk (2001a, b), Maki and Palumbo (2001), Davis and Palumbo (2001), Lettau et al (2001), and Mehra (2001).

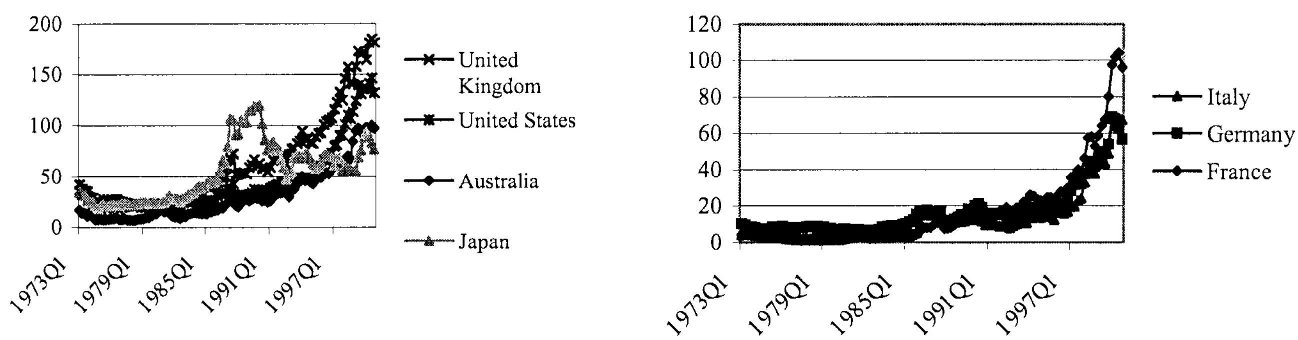
[3] See for example Girouard and Blöndal (2001), Deutsche Bank (2001) and Brady, Canner and Maki (2000).

Figure 1: Real stock price index, selected countries (1995Q1=100)



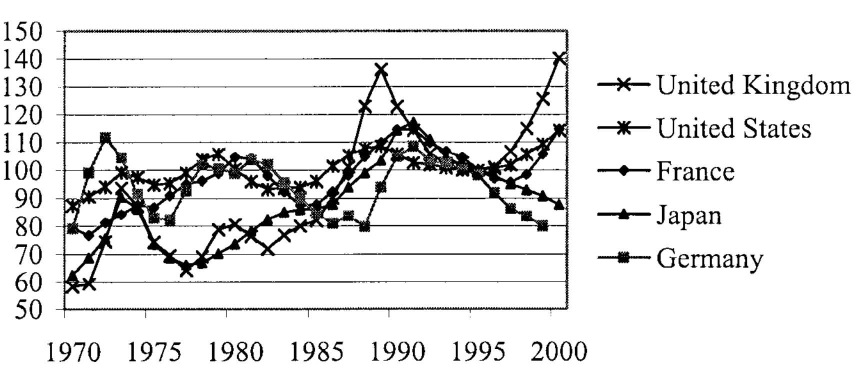
Source: International Financial Statistics, IMF (2001b)

Figure 2: Stock market capitalization as percentage of GDP



Source: Datastream

Figure 3: Real house price index (1995=100)



Source: Bank for International Settlements, (BIS, 2001) cointegrated panels developed by Pesaran, Shin, and Smith (1999) is applied in order to analyze the relationship between consumption and the two wealth components.

Most related to this contribution is a recent study by Case, Quigley, and Shiller (2001). These authors focus on the different impact of stock market and housing wealth on consumption in a panel of US states observed quarterly and a panel of 14 countries observed annually. Across various regression specifications the authors find a substantially larger impact of housing wealth on consumption than of stock market wealth. One problem associated with studies on housing wealth is the availability of housing wealth data. Case et al. impute the aggregate value of owner-occupied housing by price indices, homeownership rates, and the total number of households in a country using different data sets. This contribution relies on prices rather than actual wealth data and while presuming that prices are good proxy variables for wealth it therefore avoids assumptions on the construction of the aggregate value of housing wealth. In addition, by employing the quarterly frequency of the panel the approach adopted in this paper allows for a distinction between short run and long run impacts of changes in wealth.[4]

Previewing the results, the analysis leads to the following conclusions. First, there is a significant difference in the role of stock market wealth in the countries with market-based financial systems and countries with bank-based financial systems. In regions with market-based financial systems, the impact of both stock prices and house prices on consumption is higher in the long run. In terms of numbers, the estimated long run elasticity of consumption in stock market wealth is about twice as large in market-based economies compared with the similar coefficient for the bank-based economies. While the difference between the two groups of countries remained stable, the estimated long run elasticities have significantly increased over time mirroring not only the increased importance of stock market wealth but also the changes in financial systems. For house prices the evidence on the relative importance between the two groups is mixed and not robust against alternative specifications. However, a consistent finding is that contrary to the sample period 1960-1984, the impact of changes in housing prices on consumption was positive for the period 1985-2000 in both groups indicating that the importance of housing wealth has increased over time.

The paper is organized as follows. Section II discusses various transmission mechanisms from stock market prices and housing prices to consumption. The econometric model is introduced in Section III and the empirical results of various specifications are discussed in Section IV. Section V concludes.

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[4] Data on housing prices are only available at annual frequency and therefore different interpolation methods will be applied below.

II. Transmission Mechanisms from Wealth to Consumption

This Section analyzes the various transmission mechanisms from changes in the two variables on consumption. Tracing out the individual channels also helps to derive the main hypotheses that will be tested in the empirical investigation.

A. Channels of transmission for stock market wealth

Broadly speaking, there are five different transmission channels from changes in stock market wealth to changes in consumption[5]:

1. *Realized wealth effect*: If the value of consumers' stock holdings increase and consumers realize their gains then consumption will increase. This result will be a direct effect as a consequence of higher current liquid assets.

2. *Unrealized wealth effect*: An increase in stock prices can also have an expectations effect where the value of stocks in pension accounts and other locked-in accounts increases. When these assets go up in value but the increase is not realized, it results in higher consumption today on the expectation that income and wealth will be higher in the future.

3. *Liquidity constraints effect*: An increase in stock market prices increases the value of a portfolio for an investor. Borrowing against the value of this portfolio in turn allows the consumer to increase consumption.

4. *Stock option value effect*: An increase in stock prices can lead to higher consumption for stock option owners as a result of an increase in the value of households' stock options. Again, this increase in consumption may come independently of whether the gains are realized or unrealized.

These transmission channels are thus associated with a positive impact of changes in prices on consumption. It is worth noting the difference between the realized and unrealized wealth effect. As Poterba (2000) puts it: "it seems particularly likely that the marginal propensity to consume out of wealth gains in (locked) retirement accounts is lower than the propensity to consume out of directly held assets since the former are often thought of as 'long term assets'" (see also Thaler (1990)).[6] Furthermore, there appears to be a difference in how

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[5] Some of these channels were first identified in the life-cycle and permanent income theory (Friedman (1957), Ando and Modigliani (1963), Modigliani and Brumberg (1979)).

[6] According to Poterba and in light of the growing relative importance of retirement accounts this difference may have reduced the marginal propensity to consume out of stock market wealth in the United States. Indeed, Ludvigson and Steindel (1999) as well as Mehra (2001) estimate a lower marginal propensity to consume out of total wealth for samples of later periods.

households (and firms) adjust their consumption to changes in different types of assets. For non-technology stocks Edison and Sløk (2001a and 2001b) find a stronger reaction in consumption to changes in stock prices in market-based financial systems (United States, Canada, and United Kingdom) than in bank-based financial systems such as continental Europe. For changes in technology stock prices, however, the consumption reaction is more similar, reflecting that the technology sector worldwide seems to function and carry out business in a more similar way. Finally, as a fifth transmission channel, consumption of households who do not participate in the stock market may be indirectly affected by changes in stock market prices. Such indirect effects between stock prices and consumption have for example been highlighted in Romer (1990).

B. Channels of transmission for housing wealth

There are also five different transmission channels from changes in housing prices to changes in consumption, but some of these channels are somewhat different from the channels for stock prices listed above:

1. *Realized wealth effect*: For consumers who are house owners, the increase in house prices leads to an increase in net wealth, which can raise consumption today. If house prices increase it is possible for consumers to take out equity in the form of refinancing or selling of the house. Such a realized gain must be expected to have a positive impact on private consumption.

2. *Unrealized wealth effect*: If house prices increase but households do not refinance or sell the house it may still have a positive impact on consumption due to the increase in the discounted value of wealth. Hence consumers can spend more today on the expectation that they are "richer" than they were before.

3. *Budget constraint effect*: For consumers who are house (or apartment) renters an increase in house prices has a negative impact on private consumption. As house prices go up the budget constraint becomes tighter for renters, which must be expected to result in lower private consumption. This channel works through a realized capital loss since the increase immediately leads to higher prices, which have to be paid by the renters.[7]

4. *Liquidity constraints effect*: A fourth factor that is also important for the consumption impact of house price changes is how well functioning the financial system is. If house prices change it may require access for consumers to credit markets in order to take loans against the increase in house prices. If credit is constrained or the financial system is not

[7] This "budget constraint effect" is also but to less extend relevant for house owners since an increase in housing prices might not only increase the rent but also other expenditures on housing services such as fuel and power.

able to support such a wish for loans, households may experience that they cannot react accordingly to higher house prices.

5. *Substitution effect*: An increase in house prices may imply that households who are planning to buy a home may lower consumption when faced with increasing higher prices since they may increase down payments and future loans. This might force households to either buy a smaller house or to lower private consumption.

Both the realized and unrealized gains from increases in house prices must be expected to increase private consumption, but as with stocks, the marginal propensity to consume out of unrealized gains in housing wealth might be lower. But both the budget constraint and the substitution effect work in the opposite direction. The latter was neglected in the discussion on stock prices since the degree of divisibility is much higher for stocks and the decision to buy a house is driven by other motivations than the decision to invest in the stock market.

There are various other dimensions along which housing wealth differs from stock market wealth. For example, a similar argument to Poterba's (2000) argument about the lower marginal propensity to consume out of unrealized wealth can be made for housing wealth since it is considered as long-term asset. Moreover, the two types of assets have different risk characteristics as for example mirrored in lower loan-to-value ratios for housing wealth and there might be higher costs imbedded to getting information on housing wealth than on stock market wealth.[8]

In sum, while the effects of changes in stock market prices seem to unambiguously point to an increase in consumption, the impact of changes in housing prices is ambiguous. Further, the discussion of the various channels of transmission from movements in both wealth components to consumption suggests that the financial system plays a central role in this process.

C. The role of the financial system

The discussion above suggests that the design of the financial system plays an important role in the transmission mechanisms from changes in the two wealth components to changes in consumption. Differences in the design of the financial systems are particularly pronounced between on the one side most of the continental European countries which have bank-based financial systems and then on the other side the United States, Canada, and United Kingdom which have market-based financial systems. The design of the financial systems has important implications for the strength of the wealth effect. Of key importance for the strength are a) differences regarding sizes of financial markets and b) how widespread stock ownership is and c) the use of stock options as a means of payment by firms. These factors all influence the

[8] Other reasons for a differential impact of housing wealth and stock market wealth on consumption are discussed in Case, Quigley and Shiller (2001).

magnitude of the impact of an increase in wealth, and these different components will now be analyzed in turn.

First, the market-based system has resulted in different sizes of stock markets. In general, the size of stock markets is much higher in Anglo-Saxon countries than in the bank-based systems in Continental Europe (Figure 2). According to this criterion, the Netherlands, Japan, and Sweden might also be considered to belong to the group of market-based economies. Japan is a bit special, which has to do with significant amounts of crossholding and cross ownership of stocks. Moreover, Japan has experienced the well-known sharp increase in asset values during the eighties followed by a sharp decrease in values in the early nineties.

Second, the market-based system has also lead to a higher degree of stock market participation by households in countries with market-based financial systems than in countries with bank-based systems. Figure 4 describes some of these differences. According to this criterion, again the Netherlands and Sweden could be considered as market-based systems whereas Japan could not.

Third, stock options as a means of payment is at least until now more widespread used in Anglo-Saxon countries than in countries with bank-based financial systems. Also, numbers for venture capital clearly suggest that there is a stronger market-based tradition for venture capital in the Anglo-Saxon countries.[9]

Borio (1996) provides an analysis of the credit characteristics of the non-government sector in fourteen industrialized countries. His criteria and the discussion above suggest assigning Australia, Canada, the Netherlands, Ireland, Sweden, the United Kingdom and the United States to an extended group of market-based economies whereas Belgium, Denmark, Finland, France, Germany, Italy, Japan, Norway, and Spain are treated as bank-based.[10] The Japanese bubble economy, however, could add a lot of noise to our data and therefore the estimation was also carried out when Japan was excluded from the sample. These results do not differ from the estimates reported in section IV and are shown in tables A2 to A4 in the Appendix.

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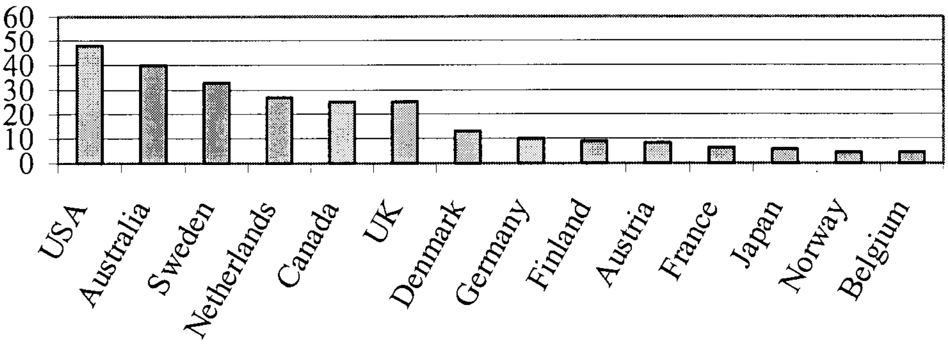
[9] See e.g. Edison and Sløk (2001) for a more general discussion of this issue.

[10] A distinction between countries according to more objective measures and formal criteria as in Beck and Levine (2001) would be warranted. See the conclusions for a more general discussion on this issue.

Figure 4. Percentage of adult population holding shares in 2000



Source: Proshare (2000)

Table 1: Homeownership in different countries

Ireland	80.0 %
Spain	78.0 %
Italy	78.0 %
United States	67.5 %
United Kingdom	67.0 %
Belgium	65.0 %
EU Average	61.0 %
Finland	60.8 %
Sweden	60.0 %
Austria	55.0 %
France	54.0 %
Denmark	52.3 %
Netherlands	52.0 %
Germany	41.0 %

Source: Hypostat (2001) and Fannie Mae Foundation (2001). Data are from the most recent year (and all data are from the 1990s).

While the distinction of these two groups of countries according to stock market capitalization and the degree of participation in the stock market is more or less distinct, such an ordering matches countries with very different homeownership rates into one group. The comparison of homeownership rates shows that three of the countries in the group of market-based economies (Ireland, the United States and the United Kingdom) have fairly high homeownership rates, whereas two others (Sweden and the Netherlands) do not (Table 1). Such an alternative grouping of the data will be considered in section IV.

Overall, these three factors support the notion that changes in stock market wealth have a larger impact on consumption in countries with market-based financial systems. At the same time, the responsiveness of consumption to changes in stock market prices must be expected to have increased substantially over time. For housing wealth, however, the impact is more uncertain. In particular, as discussed above, there is no a priori reason to expect a positive effect of changes in housing prices on consumption. But due to the deregulation of financial markets across countries and a corresponding increased information level and transparency in financial markets, it must be expected that the positive influences of housing prices on consumption have become more important over time. This effect is likely to be more pronounced in countries where households have easier access to their housing wealth, e.g. refinancing of loans. In addition, the analysis of the differences between the two wealth components suggests that consumption reacts stronger to changes in stock market wealth than to changes in housing wealth.

Based on the considerations above, Table 2 summarizes the main hypotheses to be tested in the empirical section.

Table 2: Summary of main hypotheses

Hypothesis 1	Positive effect of stock market prices on consumption
Hypothesis 2	Stock market effects are higher in countries with market-based financial systems
Hypothesis 3	The impact of housing prices on consumption is ambiguous
Hypothesis 4	Stock market effects on consumption have increased over time
Hypothesis 5	The positive effect of changes in housing prices has become more important over time
Hypothesis 6	The sensitivity of consumption to changes in stock market wealth is higher than the sensitivity to changes in housing wealth

III. THE ECONOMETRIC MODEL

In macroeconomic specifications of the consumption function, consumption is traditionally explained by income and wealth, and a main focus of this paper is the relative importance of different wealth components in different countries. A simple model of an aggregate consumption function with household (labor) income and wealth as the only determinants is motivated by several theories, including the permanent income theory by Friedman (1957) and the life cycle theory by Modigliani and Brumberg (1954) and Ando and Modigliani (1963).[11] In most empirical studies of the wealth effect of consumption, a common trend among the three variables is assumed and tested for.[12] Gali (1990) provides a theoretical foundation for a common trend approach between these three macroeconomic aggregates. In this paper it is assumed that (and tested if) such a cointegrating relationship exists between consumption, income, and the two wealth measures. Next, an error correction specification of a consumption function as first proposed in Davidson et al. (1978) is estimated for a sample of 16 OECD countries using panel data techniques.

[11] For a further discussion of the theories underlying the consumption function see Deaton (1992), Browning and Lusardi (1996), and Attanasio (1999).

[12] See e.g. the recent studies by Mehra (2001) and Lettau and Ludvigson (2001).

The literature on inference in dynamic and cointegrated panels has evolved rapidly over the past few years.[13] Among the various estimators suggested in the literature, the pooled mean group (PMG) estimator proposed by Pesaran et al (1999) is particularly attractive since it pools the long run relationship between the countries while the short run responses are flexible and unrestricted across countries. Averages of the short run coefficients across countries are computed, which is the so-called a mean group estimation (MG) in the fashion of Pesaran and Smith (1995). Therefore, this likelihood-based estimation procedure is an intermediate procedure to pooling the panel data and fully unconstrained estimation.

An identical form of the long run consumption function is assumed for all countries, where the long run relationship between consumption, income and the two wealth measures is given by

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

and the subscripts i and t denote the country and time respectively. c is the log of private per capita consumption, y^d the log of per capita disposable household income and w^{hw} and w^{sm} refer to the logs of housing and stock market wealth respectively. ε is the error term capturing the effects of unexpected shocks to consumption. Note that one may also think of the LHS of equation (1) as planned consumption (see e.g. Mehra, 2001).

Deviations from the long run relationship given by equation (1) are possible in the short run. There are various reasons for such deviations including adjustment costs, habit persistence and liquidity constraints (See Mehra (2001), Poterba (2000, p. 112) and also Campbell and Mankiw (1991)).[14] It is assumed that (short run) consumption functions differ across countries. This assumption will in the following be implemented by determining the lag length of each variable by conventional statistical criteria. For ease of presentation, it will be assumed that the first lag of each variable is an important determinant of the short run relationship in each country. The ARDL(1,1,1,1) specification of equation (1) therefore becomes

$$(2) \quad c_{t,i} = \delta_i + \beta_{10i} y^{d}_{t,i} + \beta_{11i} y^{d}_{t-1,i} + \beta_{20i} w^{sw}_{t,i} + \beta_{21i} w^{sw}_{t-1,i} + \beta_{30i} w^{hw}_{t,i} + \beta_{31i} w^{hw}_{t-1,i} + \gamma_i c_{t-1,i} + \eta_{i,t}\,.$$

The error term is assumed to be independently distributed across t and i but the variances may be heterogeneous across countries. The cross-sectional independence assumption of the error term is rather strong and restrictive. For example, it is not hard to imagine shocks that affect all countries at the same time. This assumption is standard in the dynamic panel

[13] An excellent overview of this literature is given in Baltagi and Kao (2001).

[14] Attanasio (1999) stresses the importance of adjustment costs for durable consumption, which is included in the measure of consumption considered here (see section 4).

literature and its implications for this analysis will be discussed in the conclusion. Moreover, it is assumed that the error term is independent of all the other variables in equation (2), an assumption that is invalidated if other important variables influence consumption that is not contained in equation (2). Rewriting equation (2) gives the error correction specification as:

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

where

$$(4)\qquad \phi_i = -(1-\gamma_i),\ \alpha_{0i} = \frac{\beta_i}{1-\gamma_i},\ \alpha_{1i} = \frac{\beta_{10i} + \beta_{11i}}{1-\gamma_i},\ \alpha_{2i} = \frac{\beta_{20i} + \beta_{21i}}{1-\gamma_i},\ \alpha_{3i} = \frac{\beta_{30i} + \beta_{31i}}{1-\gamma_i}.$$

In this framework, Pesaran et al. (1999) have suggested to restrict the coefficients of the error correction term in equation (3) to be equal across countries while all other short run coefficients are allowed to vary. The equation is then estimated by maximum likelihood.[15] The restriction of equal long run coefficients across countries can be tested by a conventional likelihood ratio (LR) test. As Pesaran et al (1999) point out; it is very likely that this homogeneity restriction is rejected in empirical applications. One obvious explanation for such a rejection is simple: the restriction is wrong. Another explanation is that there might be sample-specific omitted variables in the individual country regressions or measurement errors that are correlated with the regressors. While it might be possible to correct for such biases in individual country regressions, it is impossible to do so in a panel of countries. If such biases average to zero across groups, then pooling is sensible since it removes such random variation. Thus pooling might provide a more reasonable estimate of the true coefficient. If the restriction is wrong and one is interested in the average affect across a certain group of countries, then pooling and thereby ex ante imposing homogeneity might be more reasonable in small samples. While the MG estimator (the un-weighted average of individual country specific estimates) where homogeneity is imposed ex post is very sensitive to outliers in small samples, pooling which weighs the individual country specific heterogeneous coefficients according to precision reduces such bias. Along this line, the estimated coefficients can be interpreted as the weighted averages of individual group estimators while the weights are determined by the inverses of their variance covariance matrices.[16]

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[15] The approach thus differs from the conventional two-step procedure applied to single equation cointegrating relationships as in the seminal contribution of Engle and Granger (1987).

[16] More precisely, the PMG estimator can be written as a matrix weighted average of the individual group estimators while the weights are proportional to the inverses of their variance covariance matrices (see the discussion in the 1997 working paper version of Pesaran et al. (1999) on page 13).

A few further remarks on the econometric procedure are in order: First, the coefficients on the lagged dependent variables in equation (2) are subject to the familiar small sample (small T) downward bias. Since this downward bias is in the same direction for each group, averaging or pooling does not remove the bias. Second, as Pesaran and Smith (1995) points out, falsely imposing homogeneity in panels leads to an upward bias in the estimates of the coefficients on the lagged dependent variables, a bias that is not reduced when both T and N grow large. It is possible to determine if such an upward bias is serious. Under slope homogeneity, the PMG estimators are consistent and efficient while the MG estimator is consistent but inefficient. Therefore a Hausman-type test for comparison of the MG and the PMG estimators can be applied. Thus there are two biases pointing into opposite directions. However, it is not clear to what degree one bias offsets the other.

IV. Empirical Results

A. Data

Below, equation (3) will be estimated for the entire sample as well as for groups of countries according to the discussion in Section II. Data covers 16 OECD countries. Specifically, Belgium, Denmark, Finland, France, Germany, Italy, Japan, Norway and Spain will be included into the group of economies with bank-based financial systems while Australia, Canada, Ireland, the Netherlands, Sweden, the United Kingdom and the United States are treated as market-based economies. Data availability of the unbalanced panel is shown in Table A1 of the appendix.

Given the broad coverage of the study there are certain data limitations. Therefore the approach adopted below deviates a bit from the simple theoretical model formulated above. First, stock market and housing prices are here used as proxy variables for the wealth components. Note, however, that a high correlation traditionally has been found between stock market prices and wealth measures as documented in Lettau and Ludvigson (2001) and Deutsche Bank (2001). The (in)direct impact of stock market prices on aggregate consumption has for example been investigated in the studies by Romer (1990) and more recently in Poterba and Samwick (1995). The role of housing prices on consumption is the focus in, among others, Miles (1992), Miles (1995, chapter 4) and more recently Brady et al. (2000) and Girouard and Blöndal (2001).

To underscore the validity of using price data as proxy variables, the analysis is extended in section IV.F by using market capitalization as a more direct measure of stock market wealth. But even with the use of market capitalization data as a proxy for stock market wealth of households, one problem remains: international capital mobility. However, still a high proportion of stocks within each country are held by domestic residents and therefore the use of such variables as proxies for domestic stock market wealth seems to be a valid approximation.

Second, focus is on total aggregate consumption and there is no distinction between non-durable and durable consumption. Conventional theories on consumption apply to the flow of consumption. Since durable consumption can be thought of as a replacement and addition to a capital stock, the conventional approach is to only use non-durable consumption in wealth effect studies.[17] However, and as pointed out in Mehra (2001) total consumption is the parameter of interest when studying movements in stock market prices. Particularly stock market crashes are more likely to lead to a postponement of durable consumption while the reduction of non-durable consumption might be of minor importance (see e.g. Romer (1990)). Regarding the role of housing prices on aggregate consumption, durable consumption goods are among the major entities on which resources raised by mortgage refinancing are spend as Brady, Canner and Maki (2000) show. One shortcoming with using total consumption is that it also includes expenditures on housing services (see section II). Besides, it might be insightful to nevertheless distinguish between the two components of consumption but for reasons of data availability this is difficult for a panel of countries.

Third, total disposable income is used and not only labor income as would be suggested by the traditional permanent income hypothesis. Data availability constrains us to do so. But also economically it is more sensible to use total income rather than labor income. This is suggested by an extended view of the life-cycle theory as suggested by Attanasio (1999), and also sensible if households are on average more myopic than the life-cycle theory of consumption would suggests (see Campbell and Mankiw (1991) and more recently Mankiw (2000)).

Data for consumption and disposable household income was taken from the OECD Analytical Database (OECD, 2001b). Data on stock market price indices are taken from International Financial Statistics (IMF, 2001b), which provides us with a relatively broad coverage and enough time series observations. Data on housing price indices are taken from the Bank for International Settlement's house price database (BIS, 2001).[18] The data on housing prices are in annual frequency, and in order to interpolate the data linear interpolation was applied. Below this interpolation method is tested in order to check the robustness of the estimated coefficients, and it turns out that the interpolation method is in general not important for the main results found.

All variables are in local currencies and deflated by the consumer price index taken from the OECD Analytical Database. Consumption and income are expressed in per capita units using United Nations population data (UN, 2000), which are linearly interpolated between annual observations. Logs have been taken of all variables and hence the estimates reported below are the estimated elasticities of consumption in changes of the right hand side variables. We therefore control for the different size of movements in stock market and house prices across countries.

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[17] See Lettau and Ludvigson (1999, 2001) for a discussion.

[18] The comparability of these indices across countries is discussed in Girouard and Blöndal (2001, p. 36).

As Engle and Granger (1987) have pointed out, the long-run elasticities in equation (3) cannot be consistently estimated if all the single variables have unit roots unless the variables in the long-run relationship are cointegrated. Therefore, one has to examine the statistical properties of the data and test whether a cointegrating equilibrium relationship between consumption, income and the two price indices exists. Recently, tests for unit roots of individual series and cointegrating relationships between series have been developed for panel data.[19]

B. Unit root tests

Among the various tests proposed in the literature, the Im, Pesaran and Shin (1997) (IPS) panel unit root test is suitable here.[20] The IPS t-bar test is based on an average of individual country augmented Dickey Fuller (ADF) tests while allowing for heterogeneous coefficients under the alternative hypothesis and different serial correlation patterns across groups. Under the null hypothesis all groups exhibit a unit root while under the alternative this is not the case for some i.[21] A more detailed discussion of the test can be found in Baltagi and Kao (2000).

More specifically, the following model is tested for all variables.

$$\Delta y_{t,i} = \theta_i + \rho_i y_{t-1,i} + \sum_{j=1}^{p_i} \varphi_{ij} \Delta x_{i,t-j} + \upsilon_{i,t} \quad (5)$$

Under the null hypothesis the autocorrelation coefficient ρ_i equals one for all i while under the alternative it does not for at least one i. The test statistic is then computed as the average of the individual ADF statistics as

$$\bar{t} = \sum_{i=1}^{N} t_{\rho_i} \quad (6)$$

Im, Pesaran and Shin (1997) show that this test statistic converges to a standard normal distribution. Under the alternative hypothesis, the IPS panel t-bar test diverges to minus infinity and therefore the left tail of the standard normal distribution is used to reject the null hypothesis. Table 3 summarizes the results for the unit root tests of the four variables where the lag length p_i

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[19] Baltagi and Kao (2000) provide a review.

[20] The Chiang and Kao (2001) NPT 1.2 program is used to implement these tests.

[21] Choi (2001) has proposed various tests that relax upon this restrictive formulation of the alternative hypothesis in allowing for tests against the alternative where some groups exhibit a unit root and others do not.

in equation (5) is chosen by the Schwartz Bayesian criterion (SBC) with a maximum number of four lags.[22]

Table 3: IPS (1997) panel unit root tests

Variable	Test Statistic
Log of private consumption (per capita)	-1.59983
Log of household disposable income (per capita)	3.36438
Log of IFSS stock market price indices	0.43943
Log of housing price indices	-2.23799**

Note: * (**) indicate significance at the 10 (5) percent level. Test results are for the sample period of 1960-2000.

These results confirm that the null of a panel unit root is not rejected for most of the series. At a five percent level of significance it is rejected for the housing price data. Note that this might result from the linear interpolation of annual housing price data. Not surprisingly, the test statistic is -0.75741 when a time trend is included in equation (5).

C. Cointegration tests

Pedroni's (1999) tests for cointegration are used to test for the null of no cointegration in the panel of 16 OECD countries.[23] Pedroni's tests allow for a considerable degree of heterogeneity between groups with regard to the intercept, the error structure, and the cointegrating relationship. Specifically, Pedroni's tests are residual-based tests on a cointegrating regression of the following type:

$$(7) \qquad y_{i,t} = \phi_i + \sum_{j=1}^{k} \lambda_{ji,t} x_{ji,t} + \zeta_{i,t} \, .$$

The seven test statistics presented in Pedroni (1999) can be grouped into two types of statistics. The first type of statistics is based on pooling along the within-dimension of the panel; the second is based on pooling along the between-dimension (see Pedroni, 1999, p. 657). The former is constructed by first summing the numerator and the denominator of the statistics over the N groups dimension separately and then dividing while the later are average statistics in that they are constructed by first dividing the numerator and denominator and then summing over the N group dimension. This feature results in a more flexible correlation pattern of the residuals

[22] We also experimented with the Akaike information criterion which showed that the results are not sensitive to the lag selection procedure. Also, we do not report individual country specific ADF statistics here. These results are available from the authors upon request.

[23] We are grateful to Peter Pedroni for providing us with his program written in RATS.

from equation (7) across i and as discussed in Baltagi and Kao (2000) allows for an easier interpretation of the statistics if the null is rejected for the second type of statistics.[24]

Therefore, Table 4 only summarizes results for the second type of Pedroni panel cointegration tests for alternative choices of cointegrating x variables.[25] Under the null hypothesis all statistics asymptotically converge to a standard normal distribution. Under the alternative hypothesis, the statistics diverge to negative infinity. Therefore, the left tail of the normal distribution is used for a rejection of the null.

When estimating the specification in equation (7) it is first investigated whether a cointegrating relationship already exists among the *x's* on average across countries and then the specification of interest is tested. The results reported in Table 4 confirm a cointegrating relationship between the variables in the specification of interest while the cointegrating relationship between the independent variables is rejected.

Table 4: Pedroni (1999) panel cointegration tests

Regression Specification	**Group ρ-Statistic**	**Group t-Statistic (non-parametric)**	**Group adf-Statistic**
y: Income (Y^d) x_1: Stock Market Price Index (*SMPI*) x_2: Housing Price Index (*HPI*)	3.58687	2.28910	-0.23974
y: Consumption (*C*) x_1: Income (Y^d) x_2: Stock Market Price Index (*SMPI*) x_3: Housing Price Index (*HPI*)	-4.21876**	-5.24211**	-4.63842**

Note: * (**) indicate significance at the 10 (5) percent level. Test results are for the sample period of 1960-2000.

D. Estimating consumption equations

The investigation of the data properties above imply that estimation of equation (3) with variables expressed in log levels provides reliable inferences about the long and short term influences of the stock market and housing price indices and income on consumption. The considerations in Section II suggest that the impact of stock market and housing prices on consumption has changed over time. To test this hypothesis, equation (3) was estimated first for

[24] For a more detailed discussion on the statistics the reader is referred to Pedroni (1999) and Baltagi and Kao (2000).

[25] Results for the first type of tests support the findings in table 4.

the sub-period 1985Q1 to 2000Q4.[26] Regression results for the combined sample and for the two groups, market-based and bank-based economies are summarized in Table 5. In the table *C* is consumption, Y^d is disposable income, *SMPI* is the stock market price index, and *HPI* is the house price index. The Schwartz criterion is applied for selection of the lag length for each individual country with a maximum number of four lags is allowed. Given the different lag lengths applied, the short run coefficient estimates are not representative for all countries.

Table 5 shows a relatively low estimate of the income elasticity. An income elasticity less than one is suggested by economic theory in a life-cycle model inter alia by Ando and Modigliani (1963) and Gali (1990, p. 439).[27] In a regression with income as the only dependent variable, a value close to (and insignificantly different from) one was found for the entire group sample. This suggests that the presence of the wealth measures in the specification takes out some of the co-variation of income and consumption.

For the combined sample, both the estimated house price elasticity and the stock market price elasticity are positive and significant while the size of the coefficient estimate on house prices is about half the size of the coefficient estimate on stock market prices. The difference is significant with a t-ratio of 3.28. Splitting the sample into the two groups reveals that the estimated coefficients are roughly similar for the market-based economies while they are substantially reduced for the bank-based economies. This similarity of the coefficient estimates for the market-based economies to the estimates of the combined sample of all countries indicates that the market-based economies have a higher relative weight than the bank-based economies in the combined sample.[28] For stock market prices, the difference between the estimated elasticities for the two groups of countries (and its t-ratio) is 0.051 (4.89). The estimated housing price elasticity for the bank-based economies is insignificant and also lower than the significant estimate for the market based economies but the difference in the estimates between the two groups of countries is insignificant.

Thus, the estimated coefficients on stock market wealth are about twice as large as the estimated coefficients on housing wealth. This finding is consistent with our earlier analysis on the differences between the two wealth components (Section II). Yet it contrasts with the study of Case et al. (2001) who find a remarkably strong sensitivity of consumption to changes in housing wealth across countries (11 to 17 percent) whereas their estimated elasticity of consumption to changes in stock market wealth is rather low and unstable across different

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[26] Effectively, this time period is shorter and differs from country to country. The shortest time period is 42 observations for Belgium and the longest are 59 observations for Australia, Canada, France, Germany and the United States.

[27] Its exact value depends among other things on the age structure of population and the distribution of income.

[28] See the discussion of the PMG estimator in section III.

econometric approaches. In five out of six specifications in Case et al. the hypothesis that housing wealth has a stronger impact on consumption than stock market wealth is accepted. With regard to the different approaches between the two studies discussed in the introduction this difference in evidence must be interpreted with care. Yet Case et al. do not provide any explanation for their finding and given that housing wealth is supposedly less liquid, viewed more as a long term asset and that information on its exact value is not as widespread as for the value of stock market wealth, a higher sensitivity of consumption to changes in housing wealth seems counterintuitive and might have resulted from the way housing wealth was constructed in their data set.

The size of the adjustment coefficient is higher for the two separate samples than for the combined sample of all countries. While the individual group estimates of the adjustment coefficient remain unaltered for the market-based economies they increase for four countries (Finland, France, Norway and Spain) among the group of bank-based economies. Moreover, the mean group estimate of the adjustment coefficient is higher for the bank-based economies. The difference is insignificant but nevertheless prior considerations would have suggested the opposite. However, the estimated adjustment coefficients are positive for Australia and Canada and taking the average of the remaining five countries results in an average adjustment coefficient of –0.18. Both observations illustrate the sensitivity of mean group estimation to outliers in the small sample. Table 5 also shows that the mean group estimator on the first short run income coefficient is higher for the market-based economies. This result was expected and is insensitive to outliers.

Table 5 does not show any measure for the goodness of fit of the regression specifications since achieving a best possible fit is not a primary issue of this investigation. However the un-weighted average R^2 of the individual restricted country regressions is around 50 percent for all the regressions. The individual estimates of the fit vary a lot ranging from around 21 percent in the case of Canada to 91 percent in the case of Belgium (and this pattern again is independent of the estimated specification).

The likelihood ratio statistic that tests for the homogeneity assumption of the long run coefficients is rejected for all groups. A comparison of the sum of the two independent likelihood statistics for the two groups with the statistic for the combined panel reveals a modest gain from grouping the data into the two groups. However, equality of the mean group and pooled mean group estimators is accepted by the Hausman test for all groupings of the data. This means that the pooled estimates are not biased by the imposition of homogeneity.

Table 5: Estimating consumption functions (1985-2000)

Variables	All Countries	Bank-based Economies	Market-based Economies
		Long run coefficients	
Yd	0.7031** (0.0227)	0.6444** (0.0471)	0.7056** (0.0275)
SMPI	0.0802** (0.0059)	0.0305** (0.0077)	0.0815** (0.0072)
HPI	0.0362** (0.0122)	0.0154 (0.0186)	0.0403** (0.0215)
		Averages of heterogeneous short run coefficients	
Adjustment coefficient	-0.096** (0.062)	-0.140* (0.123)	-0.128** (0.043)
d(C(-1))	0.006 (0.1)	-0.016 (0.173)	0.099 (0.109)
d(Y^d)	0.276** (0.083)	0.346** (0.138)	0.205** (0.085)
d($Y^d(-1)$)	-0.075 (0.080)	-0.109 (0.141)	-0.058 (0.090)
d(*SMPI*)	-0.003 (0.002)	-0.005 (0.004)	-0.003 (0.003)
d(*SMPI(-1)*)	-0.002 (0.002)	0.000 (0.000)	-0.006 (0.006)
d(*HPI*)	0.162** (0.057)	0.142* (0.087)	0.180** (0.086)
d(*HPI(-1)*)	-0.111** (0.057)	-0.084 (0.068)	-0.108 (0.108)
Intercept	0.063** (0.019)	0.209** (0.070)	0.029** (0.014)
Diagnostic Statistics			
LR test (p-value)	245.47 (0.00)	149.93 (0.00)	83.39 (0.00)
Joint Hausman test (p-value)	5.62 (0.13)	4.09 (0.25)	7.67 (0.05)

Note: See Table A1 for a description of the variables. The LR test tests for the homogeneity restriction of the long run coefficients. The Hausman test tests for equality of the MG and the PMG estimators. The Hausman test statistic is indeterminate if the difference between the variance-covariance matrices of the MG and PMG estimators is not positive definite (see Pesaran et al. (1999) for more details). The unrestricted short run coefficient estimates are the MG estimates under the restriction of long run homogeneity. Standard errors of the estimated coefficients are in parenthesis. * (**) indicate significance at the 10 (5) percent level.

In sum, the evidence supports the hypothesis of a positive responsiveness of consumption to changes in the two wealth components on average over the last 15 years. The impact of price changes is higher for the group of market-based economies than it is for the group of bank-based economies and consumption is more sensitive to changes in stock market wealth than housing wealth.

E. Sensitivity analysis

Next, it is investigated whether the coefficient estimates shown in Table 5 have changed across time by estimating equation (3) for the period 1960Q1 to 1984Q4. Again, the effective time period is shorter and the number of observations range from 28 in the case of Ireland to 93 for Australia. As before, the Schwarz criterion is applied for the selection of the lag length for each country regression. Estimation results are shown in Table 6.

The results confirm a number of the hypotheses discussed above. First, the estimated stock market price has increased both for the entire sample as well as for both groups. In particular, the stock market elasticity has been very low and insignificant for the bank-based economies during the first observation period. As for the second observation period, the estimated elasticity is significantly higher for the market-based economies. Second, the estimates of the housing price elasticities are all negative. This suggests that during the period from 1960 to 1984 the negative impact of housing prices on consumption dominated the positive impact. The change in sign of the housing price coefficient is a striking evidence for chances in financial markets, particularly in the mortgage markets during the late 1980's, which has made it easier for households to access their housing wealth (see Brady, Canner and Maki (2000) and Girouard and Blöndal (2001)).

The specification of the consumption function used here leaves out some variables that also influence consumption. Most importantly, only rather narrow measures of wealth were used. Including other independent variables like the short interest rate, government consumption and measures of the population age distribution either as exogenous regressors or in the long run relationship had only minor effects on the results reported above. However, Lettau, Ludvigson, and Barczi (2001) critically discuss the inclusion of such variables as exogenous regressors in a cointegrating framework like above since this results in a biased adjustment coefficient in a two-step Engle-Granger procedure if these variables are not weekly exogenous. In a maximum likelihood approach as adopted here, including such variables might not only bias the adjustment coefficient but also the cointegrating relationship itself.

Beyond such robustness checks the robustness of the housing price estimates with regard to alternative measures of housing prices was investigated. In particular, it was checked whether the earlier estimates are robust with regard to alternative interpolation strategies like cubic spline. Also for a small sub-sample of countries regressions were run with actual quarterly

Table 6: Estimating consumption functions (1960-1984)

Variables	**All Countries**	**Bank-based Economies**	**Market-based Economies**
		Long run coefficients	
Y^d	0.9171** (0.0159)	0.9045** (0.0191)	0.9144** (0.0273)
SMPI	0.0152** (0.0045)	0.0059 (0.0057)	0.0263** (0.0079)
HPI	-0.0544** (0.0131)	-0.0377** (0.0185)	-0.0282 (0.0230)
		Averages of heterogeneous short run coefficients	
Adjustment coefficient	-0.213** (0.051)	-0.218* (0.074)	-0.223** (0.078)
d(C(-1))	0.130 (0.073)	0.202 (0.124)	0.039 (0.046)
d(Y^d)	0.355** (0.077)	0.386** (0.105)	0.306** (0.115)
d(Y^d(-1))	-0.095 (0.067)	-0.176 (0.117)	0.006 (0.006)
d(*SMPI*)	0.012** (0.005)	0.007 (0.006)	0.018** (0.008)
d(*SMPI(-1)*)	0.0001 (0.003)	0.00 (0.00)	0.001 (0.007)
d(*HPI*)	0.009 (0.014)	-0.006 (0.006)	0.022 (0.030)
d(*HPI(-1)*)	0.00 (0.00)	0.00 (0.00)	0.00 (0.00)
Intercept	0.094** (0.026)	0.116** (0.041)	0.043* (0.022)
Diagnostic Statistics			
LR test (p-value)	188.95 (0.00)	86.33 (0.00)	99.5588 (0.00)
Joint Hausman test (p-value)	n.a.	n.a.	n.a.

Note: See Table A1 for a description of the variables. The LR test tests for the homogeneity restriction of the long run coefficients. The Hausman test tests for equality of the MG and the PMG estimators. The Hausman test statistic is indeterminate if the difference between the variance-covariance matrices of the MG and PMG estimators is not positive definite (see Pesaran et al. (1999) for more details). The unrestricted short run coefficient estimates are the MG estimates under the restriction of long run homogeneity. Standard errors of the estimated coefficients are in parenthesis. * (**) indicate significance at the 10 (5) percent level.

housing price data that were provided by the ECB.[29] For this small set of countries (Germany, Italy, Spain, Sweden and the United Kingdom) and generally shorter time series, the estimates confirmed the positive influence of changes in housing prices on consumption that was found for the second observation period. An alternative grouping of the data according to homeownership rates as discussed in section II found a negative estimate for the long run elasticity of consumption to housing prices for the first period in both groups. In the second period it was, however, positive for the group of countries with relatively high ownership rates and negative for the other group, supporting our earlier findings of important interactions between financial markets and ownership of assets. Further, the robustness of the stock market price elasticities with data on stock market capitalization as a more direct measure of stock market wealth was investigated. The next subsection will turn to an analysis of these results.

F. From elasticities to marginal propensities to consume

In this subsection the elasticities estimated above are translated into estimates of marginal propensities to consume out of stock market wealth. For this, the regressions shown in Table 5 were estimated using logs of real stock market capitalization per capita instead of the price data. The stock market capitalization data are in quarterly frequency and taken from Datastream. Since comparable data on housing wealth are not available it is impossible to perform the same exercise for housing wealth.

It is compelling how close the estimates for the long run elasticities to consume out of stock market wealth are to the earlier estimates in Table 5. The additional quantity dimension of the market capitalization data does not affect these estimates but changes the coefficient estimates of the long run income elasticities and the housing price coefficient, which is now higher for the bank-based economies (the difference is significant with a t-ratio of 4.3). The size of the estimated coefficient of housing wealth for the bank based economies is at the lower end of the range of estimates found in Case et al. (2001). However, compared to the other estimates it is unsystematically high and probably due to bias induced by the measurement error resulting from the interpolation of the housing price data. The estimated adjustment coefficient is now higher for the market-based economies while the difference between the two estimates is insignificant (t-ratio of 0.43). This is due to the fact that Australia and Canada now show negative coefficients and thus more reasonable coefficient estimates.

In order to translate the estimated elasticities into long run marginal propensities to consume (MPC) out of wealth, the estimates are multiplied with the recent aggregate consumption to stock market capitalization ratio. The short run estimates are accordingly obtained by multiplying the long run estimates with the estimated adjustment coefficient.[30]

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[29] We are grateful to Matteo Iacoviello (European Central Bank) for providing us with the data. A more detailed description of the data can be found in Iacoviello (2001).

[30] The coefficient estimates on the lagged and first differenced variables are small in magnitude and not representative for all countries and therefore neglected.

Table 7: Estimating consumption functions – Using stock market capitalization data (1985-2000)

Variables	**All Countries**	**Bank-based Economies**	**Market-based Economies**
		Long run coefficients	
Y^d	0.4759** (0.0218)	0.4551** (0.0276)	0.4690** (0.0232)
SMPI	0.0950** (0.0029)	0.0382** (0.0032)	0.0963** (0.0031)
HPI	0.0353** (0.0133)	0.1069** (0.0096)	0.0309** (0.0145)
		Averages of heterogeneous short run coefficients	
Adjustment coefficient	-0.131** (0.035)	-0.175* (0.065)	-0.221** (0.060)
d(C(-1))	0.067 (0.099)	0.061 (0.152)	0.114 (0.114)
d(Y^d)	0.256** (0.085)	0.325** (0.126)	0.153* (0.087)
d($Y^d(-1)$)	-0.108 (0.077)	-0.148 (0.110)	-0.048 (0.092)
d(*SMCAP*)	-0.005 (0.002)	-0.003 (0.003)	-0.009* (0.005)
d(*SMCAP(-1)*)	-0.002 (0.002)	0.000 (0.000)	-0.005 (0.005)
d(*HPI*)	0.147** (0.040)	0.173** (0.074)	0.203** (0.064)
d(*HPI(-1)*)	-0.081 (0.054)	-0.045 (0.052)	-0.111 (0.111)
Intercept	0.086** (0.026)	0.282** (0.105)	0.074** (0.023)
Diagnostic Statistics			
LR test (p-value)	236.73 (0.00)	156.44 (0.00)	50.09 (0.00)
Joint Hausman test (p-value)	14.90 (0.00)	3.49 (0.32)	n.a.

Note: See Table A1 for a description of the variables. The LR test tests for the homogeneity restriction of the long run coefficients. The Hausman test tests for equality of the MG and the PMG estimators. The Hausman test statistic is indeterminate if the difference between the variance-covariance matrices of the MG and PMG estimators is not positive definite (see Pesaran et al. (1999) for more details). The unrestricted short run coefficient estimates are the MG estimates under the restriction of long run homogeneity. Standard errors of the estimated coefficients are in parenthesis. * (**) indicate significance at the 10 (5) percent level.

Table 9 reports such estimates for a group of countries. The estimated long run MPC for the United States is 0.04. A dollar increase in stock market capitalization therefore causes consumption to increase by 4 cents in the long run which translates into a short run adjustment of about 0.8 cents increase in consumption per quarter. These estimates are in the range of estimates conventionally suggested in the literature.

For comparability with the estimate of the United States' MPC, the estimates of the MPCs of all other countries are adjusted by the variation in the stock market capitalization to GDP ratio of each particular country relative to the United States. Note that this is just another way to express the estimated elasticities from before. The average of the adjusted long run MPCs is 4.3 cents for the market-based economies and 2.6 cents for the bank-based economies.

Table 8: Marginal propensities to consume out of stock market wealth

	AU	CA	UK	US	JAP	FR	GE	IT
Adj. L.R. MPC	0.043	0.040	0.049	0.040	0.040	0.014	0.020	0.030
Adj. S.R. MPC	0.0094	0.0089	0.0109	0.0089	0.007	0.0025	0.0035	0.0053

Note: The marginal propensities to consume out of stock market wealth are calculated by multiplying the estimated elasticities with the most recent consumption to wealth ratios. Stock market wealth in each country is measured by stock market capitalization data. The adjusted marginal propensities to consume are calculated by correcting the original marginal propensity of country *i* by the variation in the first differenced series of the stock market capitalization to GDP ratio of country *i* relative to the US.

V. Conclusions and Policy Recommendations

The analysis in this paper leads to five conclusions. First, estimating panels using different groups of OECD countries shows that there is a significant long-run impact from stock market wealth to private consumption. Second, there is 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. According to our estimates, the average half time of the adjustment process to restore the long-run equilibrium is approximately 5 quarters among the countries considered. Third, there is clear evidence that the impact from changes in stock prices on consumption is bigger in economies with market-based financial systems than in economies with bank-based financial systems. Fourth, this impact from stock markets to consumption has increased over time for both countries with market-based financial systems and countries with bank-based financial systems. Fifth, and finally, while the effect of housing prices on consumption is ambiguous the wealth effect has become more important over time. For the sample period 1985-2000 the effect of housing prices on consumption is significantly positive. The estimated elasticity of house prices on consumption is about twice as large as the elasticity of stock market prices for the combined sample of all countries and for the group of market-based economies.

Before drawing policy conclusions from these results, the limitations of the employed approach have to be discussed and related to the existing literature. First, it would be preferable to have a more "endogenous" grouping of countries into bank-based and market-based rather than just grouping them by assumption. Making the determination of bank-based vs. market-based a part of the regression could for example be done by using data for outstanding mortgage loans and data for relative to private sector credit (see e.g. Beck and Levine, 2001). These series put in relation to GDP could give some indication as to whether the financial system relies more on financial markets than on banks when consumers and firms finance themselves. Along the same lines, it might be interesting to extend the analysis of the effects of deregulation of financial markets by similar methods (see the review of the literature in Boone et al., 2001).

Second, the relative difference of the estimates between the two groups of countries are roughly proportional to the differences in stock market ownership and the share of equity wealth in total wealth among OECD countries (OECD, 2001a). Since the distribution of equity wealth among the population was not controlled for, it is therefore not clear to what extend these differences in the estimates between the two groups are a mere accounting identity or related to substantial differences between the design of the financial markets in these countries. The change of the impact of housing prices suggests that financial markets play a crucial role but more research addressing this issue is needed.

Third, a direct wealth effect is not the only explanation for the observed relationship between stock market prices and consumption. Among the alternative explanations, the "leading indicators" and the "consumer confidence" channel are worth discussing. The extreme formulation of the former has been criticized, among others, by Shiller (1995): "Could it possibly be meaningful to say that there is, ultimately, no wealth effect from the stock market on consumption; that is, that if people were given more stock, they would not consume more?" The mere fact, that the coefficient estimates differ between the two groups of countries shows that there must be something else at work unless the role of the stock market as a leading indicator is significantly weaker in bank-based economies. However, this difference does not rule out the role of the "consumer confidence" channel. According to Romer (1990), a decrease in stock prices leads to an increased uncertainty about future income, i.e. a decrease in consumer confidence, and therefore to a decrease in (durable) consumption. If information about decreases in stock market prices increases with the participation of households in the stock market, then the "consumer confidence" channel is the stronger the more households participate in the stock market in a given country.

However, the difficulty in distinguishing this indirect effect from the direct wealth effect does not diminish the causal relationship from stock prices to consumption.[31] Hence it is possible to draw a number of policy conclusions. First, the stock market has become more and more important over time as a determinant of consumption. This result holds both for countries with market-based financial systems and for countries with bank-based financial systems. This has serious implications for policymakers, and in particular for monetary policy, and the results seem independent of how the financial system is designed. In other words, monetary policy should keep a close track of developments in stock prices as dramatic changes in equity values may have significant impacts on consumption, and this result seems to be relevant both in continental European countries, United Kingdom, United States, Canada, and Japan.

[31] Such a distinction would only be possible by the analysis of micro data (see e.g. Maki and Palumbo, 2001).

Second, changes in house prices also positively affect consumption and evidence is found that this effect also has become stronger over time. Most recently, the house wealth effect has functioned through the refinancing of loans due to the significant drop in mortgage rates experienced in most OECD countries.

Third, and finally, the increase in both stock prices and house prices has generated a fall in savings in particular in the United States. The strong significance of both stock prices and house price wealth in explaining consumption suggests that savings are highly correlated with developments in these two wealth components. As the *Economist* has recently argued, "the danger is that the wealth effect from housing may follow the equity wealth effect by turning negative" (Economist, 2001, p. 70). In our sample of 16 OECD countries the average correlation between the two wealth components is 0.26 on average across all countries for the entire period and has increased on average from the first to the second observation period and for thirteen out of the sixteen countries. However, as Poterba (2000, p. 110) puts it "the link between stock prices and real estate is [...] sketchy" and differs across regions.[32] A further empirical investigation of recent developments in the link between the two wealth components is certainly warranted.

Finally, some remarks on the econometric model are in order. One crucial assumption of the panel cointegration literature is the independence assumption of the error term. The high integration of national markets makes it very likely that this assumption is invalidated. Pesaran et al. (1999) and Banerjee et al. (2000) suggest various cures to this problem but since the estimates reported above are "sensible" the invalidation of this assumption seems of minor importance. A more rewarding alternative to the econometric technique might be a panel VAR approach to such a cointegration framework like suggested in Larsson et al. (2001) and Larsson and Lyhagen (1999). Adopting a higher dimensional framework to the questions addressed in this paper would allow further insights into the potentially different role between the two groups of countries that the other variables play in restoring the long run equilibrium between consumption, income and household wealth.

[32] For more detailed analysis see the literature cited in Poterba (2000).

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Table A1: Data availability

Country/Variable	*C*	Y^d	*SMPI*	*SMC*	*HPI*	*CPI*	*Population*
AU	1960Q1:2000Q4	1960Q1:2000Q4	1960Q1:2000Q4	1973Q1:2000Q4	1960:2000	1960Q1:2000Q4	1960:1999
BE	1960Q1:1999Q4	1970Q4:1999Q4	1960Q1:1996Q2	1973Q1:2000Q4	1960:2000	1960Q1:2000Q4	1960:1998
CA	1961Q1:2000Q4	1961Q1:2000Q4	1960Q1:2000Q4	1973Q1:2000Q4	1960:2000	1960Q1:2000Q4	1960:2000
DK	1960Q1:2000Q4	1960Q1:1999Q4	1960Q1:2000Q4	1973Q1:2000Q4	1960:2000	1970Q1:2000Q3	1960:2000
FI	1960Q1:2000Q4	1970Q1:2000Q4	1960Q1:2000Q4	1988Q2:2000Q4	1970:2000	1970Q1:2000Q4	1960:2000
FR	1963Q1:2000Q4	1970Q1:2000Q3	1960Q1:2000Q4	1973Q1:2000Q4	1960:2000	1960Q1:2000Q4	1960:2000
GE	1960Q1:2000Q4	1960Q1:2000Q4	1970Q1:2000Q4	1973Q1:2000Q4	1960:2000	1961Q1:2000Q4	1960:2000
IE	1960Q1:1999Q4	1960Q1:1998Q4	1960Q1:2000Q4	1973Q1:2000Q4	1960:2000	1960Q1:2000Q3	1960:1999
IT	1960Q1:2000Q4	1960Q1:1996Q4	1960Q1:2000Q4	1973Q1:2000Q4	1960:2000	1960Q1:2000Q4	1960:1998
JAP	1960Q1:2000Q4	1960Q1:1999Q4	1960Q1:2000Q4	1973Q1:2000Q4	1960:2000	1960Q1:2000Q4	1960:1999
NE	1960Q1:2000Q4	1960Q1:1999Q4	1960Q1:2000Q4	1973Q1:2000Q4	1960:2000	1960Q2:2000Q4	1960:1999
NO	1960Q1:2000Q4	1975Q1:1999Q4	1960Q1:2000Q4	1980Q1:2000Q4	1970:2000	1970Q1:2000Q4	1960:1999
SP	1960Q1:2000Q4	1964Q1:1999Q4	1961Q1:2000Q4	1987Q2:2000Q4	1975:2000	1975Q1:2000Q4	1960:1999
SWE	1960Q1:2000Q4	1960Q1:2000Q4	1960Q1:2000Q4	1982Q1:2000Q4	1970:2000	1970Q1:2000Q4	1960:1999
UK	1960Q1:2000Q4	1960Q1:2000Q4	1960Q1:1999Q1	1973Q1:2000Q4	1960:2000	1960Q1:2000Q4	1960:1998
US	1960Q1:2000Q4	1960Q1:2000Q4	1960Q1:2000Q4	1973Q1:2000Q4	1960:2000	1960Q1:2000Q4	1960:2000

Sources: OECD Analytical Database, International Financial Statistics, Bank for International Settlements and Datastream.

Table A2: Estimating consumption functions (1985-2000) – excluding Japan

Variables	**All Countries**	**Bank-based Economies**	**Market-based Economies**
		Long run coefficients	
Y^d	0.7009** (0.0227)	0.6398** (0.0486)	0.7056** (0.0275)
SMPI	0.0804** (0.0059)	0.0309** (0.0079)	0.0815** (0.0072)
HPI	0.0366** (0.0124)	0.0117 (0.0194)	0.0403** (0.0215)
		Averages of heterogeneous short run coefficients	
Adjustment coefficient	-0.094** (0.027)	-0.138* (0.062)	-0.128** (0.043)
d(C(-1))	0.046 (0.098)	0.049 (0.182)	0.099 (0.109)
d(Y^d)	0.294** (0.087)	0.390** (0.149)	0.205** (0.085)
d($Y^d(-1)$)	-0.080 (0.085)	-0.122 (0.160)	-0.058 (0.090)
d(*SMPI*)	-0.004 (0.002)	-0.006 (0.004)	-0.003 (0.003)
d(*SMPI(-1)*)	-0.003 (0.003)	0.000 (0.000)	-0.006 (0.006)
d(*HPI*)	0.173** (0.06)	0.158* (0.097)	0.180** (0.086)
d(*HPI(-1)*)	-0.118** (0.061)	-0.094 (0.076)	-0.108 (0.108)
Intercept	0.054** (0.017)	0.196** (0.080)	0.029** (0.014)
Diagnostic Statistics			
LR test (p-value)	237.54 (0.00)	138.65 (0.00)	83.39 (0.00)
Joint Hausman test (p-value)	8.29 (0.04)	2.53 (0.47)	7.67 (0.05)

Note: See Table A1 for a description of the variables. The LR test tests for the homogeneity restriction of the long run coefficients. The Hausman test tests for equality of the MG and the PMG estimators. The Hausman test statistic is indeterminate if the difference between the variance-covariance matrices of the MG and PMG estimators is not positive definite (see Pesaran et al. (1999) for more details). The unrestricted short run coefficient estimates are the MG estimates under the restriction of long run homogeneity. Standard errors of the estimated coefficients are in parenthesis. * (**) indicate significance at the 10 (5) percent level.

Table A3: Estimating consumption functions (1960-1984) – excluding Japan

Variables	All Countries	Bank-based Economies	Market-based Economies
		Long run coefficients	
Y^d	0.9240** (0.0165)	0.9138** (0.0200)	0.9144** (0.0273)
SMPI	0.0161** (0.0045)	0.0028 (0.005)	0.0263** (0.0079)
HPI	-0.0487** (0.0131)	-0.0194 (0.0167)	-0.0282 (0.0230)
		Averages of heterogeneous short run coefficients	
Adjustment coefficient	-0.223** (0.054)	-0.243** (0.083)	-0.223** (0.078)
d(C(-1))	0.138 (0.078)	0.227 (0.137)	0.039 (0.046)
d(Y^d)	0.357** (0.082)	0.393** (0.118)	0.306** (0.115)
d(Y^d(-1))	-0.115 (0.068)	-0.227* (0.120)	0.006 (0.006)
d(*SMPI*)	0.009** (0.004)	0.002 (0.001)	0.018** (0.008)
d(*SMPI(-1)*)	0.0001 (0.003)	0.00 (0.00)	0.001 (0.007)
d(*HPI*)	0.009 (0.014)	-0.008 (0.008)	0.022 (0.030)
d(*HPI(-1)*)	0.00 (0.00)	0.00 (0.00)	0.00 (0.00)
Intercept	0.086** (0.024)	0.101** (0.036)	0.043* (0.022)
Diagnostic Statistics			
LR test (p-value)	183.55 (0.00)	80.03 (0.00)	99.5588 (0.00)
Joint Hausman test (p-value)	n.a.	n.a.	n.a.

Note: See Table A1 for a description of the variables. The LR test tests for the homogeneity restriction of the long run coefficients. The Hausman test tests for equality of the MG and the PMG estimators. The Hausman test statistic is indeterminate if the difference between the variance-covariance matrices of the MG and PMG estimators is not positive definite (see Pesaran et al. (1999) for more details). The unrestricted short run coefficient estimates are the MG estimates under the restriction of long run homogeneity. Standard errors of the estimated coefficients are in parenthesis. * (**) indicate significance at the 10 (5) percent level.

Table A4: Estimating consumption functions (1985-2000) – Using Stock Market Capitalization Data, Excluding Japan

Variables	All Countries	Bank-based Economies	Market-based Economies
		Long run coefficients	
Y^d	0.4740** (0.0219)	0.4531** (0.0276)	0.4690** (0.0232)
SMPI	0.0951** (0.0029)	0.0382** (0.0032)	0.0963** (0.0031)
HPI	0.0343** (0.0143)	0.1073 (0.0095)	0.0309** (0.0145)
		Averages of heterogeneous short run coefficients	
Adjustment coefficient	-0.131** (0.035)	-0.183* (0.073)	-0.221** (0.060)
d(C(-1))	0.111 (0.095)	0.139 (0.148)	0.114 (0.114)
d(Y^d)	0.274** (0.089)	0.366** (0.136)	0.153* (0.087)
d(Y^d(-1))	-0.115 (0.082)	-0.167 (0.123)	-0.048 (0.092)
d(*SMCAP*)	-0.005 (0.002)	-0.003 (0.003)	-0.009* (0.005)
d(*SMCAP(-1)*)	-0.002 (0.002)	0.000 (0.000)	-0.005 (0.005)
d(*HPI*)	0.157** (0.041)	0.195** (0.081)	0.203** (0.064)
d(*HPI(-1)*)	-0.087 (0.057)	-0.051 (0.059)	-0.111 (0.111)
Intercept	0.074** (0.024)	0.273** (0.119)	0.074** (0.023)
Diagnostic Statistics			
LR test (p-value)	227.50 (0.00)	144.92 (0.00)	50.09 (0.00)
Joint Hausman test (p-value)	22.87 (0.00)	1.96 (0.58)	n.a.

Note: See Table A1 for a description of the variables. The LR test tests for the homogeneity restriction of the long run coefficients. The Hausman test tests for equality of the MG and the PMG estimators. The Hausman test statistic is indeterminate if the difference between the variance-covariance matrices of the MG and PMG estimators is not positive definite (see Pesaran et al. (1999) for more details). The unrestricted short run coefficient estimates are the MG estimates under the restriction of long run homogeneity. Standard errors of the estimated coefficients are in parenthesis. * (**) indicate significance at the 10 (5) percent level.