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Wealth and consumption: an assessment of the international evidence

Vincent Labhard, Gabriel Sterne and Chris Young

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*Vincent Labhard**

*Gabriel Sterne***

and

Chris Young†

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* Conjunctural Assessment and Projections Division, Bank of England.
Email: vincent.labhard@bankofengland.co.uk

** Monetary Assessment and Strategy Division, Bank of England.
Email: gabriel.sterne@bankofengland.co.uk

† Office of Charles Bean, Bank of England.
Email: chris.young@bankofengland.co.uk

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Abstract

The main objective of this paper is to offer a critique of the existing literature on the link between wealth and consumption, as captured by the long-run marginal propensity to consume from financial wealth (mpc_w). The international evidence suggests that the mpc_w varies considerably across countries, and new estimates are presented, based on structural vector autoregressions (VARs) for eleven OECD countries, which tend to confirm this finding. It is argued that there is little theoretical rationale for a wide cross-country dispersion of the mpc_w, and that the cross-country differences in empirical estimates may in fact reflect difficulties in the measurement of wealth across countries and a failure to account for the shocks causing changes in both consumption and wealth. Using a suitable panel technique, it is found that the hypothesis of a common long-run mpc_w across countries cannot be rejected consistently, and a plausible estimate is obtained for the cross-section of eleven OECD countries. This estimate is a little over 6%, broadly consistent with estimates used in a wide range of policy models.

Key words: Wealth effects, consumption, dynamic panel.

JEL classification: C22, C23, E21.

Summary

Since the mid-1990s, there have been remarkable changes in stock market capitalisation in many of the major economies, related in part to changes in the valuation of equities. Between Autumn 1994 and Autumn 2000, stock market capitalisation increased by over 100% of GDP in the United Kingdom and the United States; while between 2000 Q3 and 2003 Q2, market capitalisation fell by 65% of GDP in the United States and by 87% in the United Kingdom. These changes have motivated renewed interest in the wealth effect on consumption.

The wealth effect on consumption is often captured by the marginal propensity to consume from financial wealth, (mpc_w). The existing empirical literature suggests that this quantity varies greatly across countries, and new results presented here, based on single-country structural vector autoregressions (VARs) for eleven OECD countries, tend to confirm this finding. This divergence though is at odds with the values used in calibrated models, which tend to be more similar across countries. The main objective of this paper is to offer a critique of the literature, and to assess several possible explanations that might justify such differences in the mpc_w across countries.

It is concluded that many potential explanations cannot account for the magnitude of the differences reported in the empirical literature on the mpc_w, including differences in demographics and in the type of assets held across economies. It is argued that there is little theoretical rationale for such a wide cross-country dispersion of empirical estimates. In part, this may be due to the empirical approach taken in much of the literature: partial equilibrium approaches to capturing the impact of changes in wealth on consumption face a cocktail of data problems and cannot account for underlying structural causes of simultaneous changes in both consumption and wealth. For example, in circumstances in which there are shocks to expected earnings, economies where market capitalisation is low and wealth held in unquoted equities is *under*-recorded might be (inaccurately) estimated to have a *higher* mpc_w. Because conventional empirical estimates of the mpc_w are commonly calculated by dividing an empirical estimate of the partial elasticity of consumption with respect to wealth by the observed wealth to consumption ratio, if the reaction of consumption to an earnings shocks is similar, but wealth is under-recorded because of data problems, then the mpc_w will be overestimated.

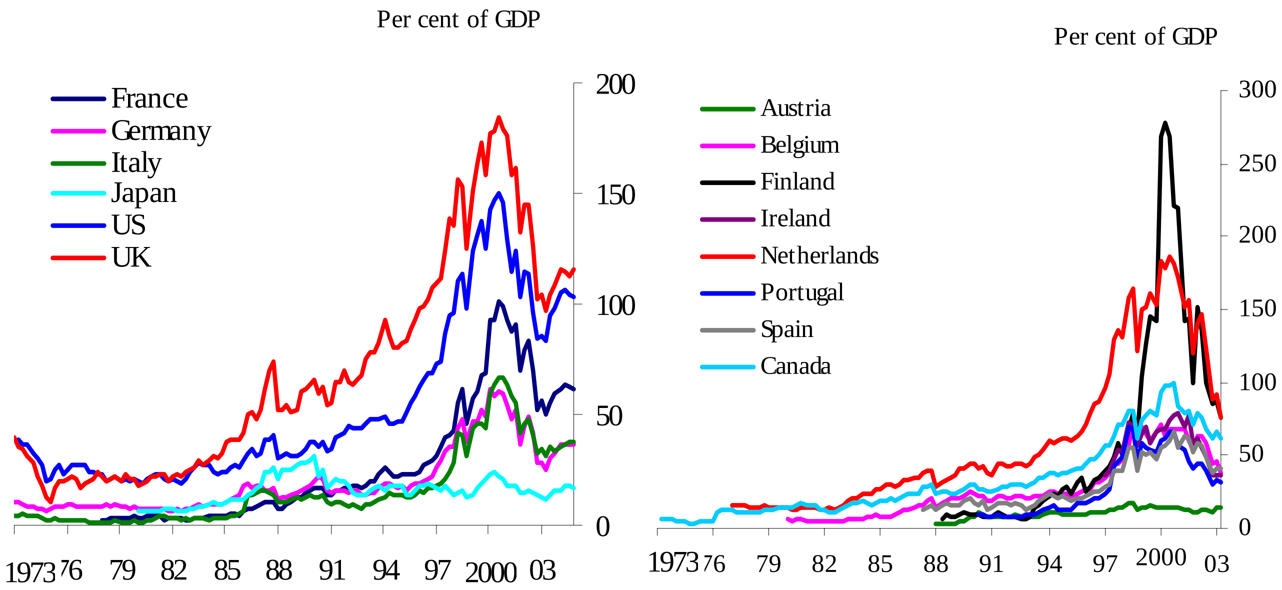
This leaves the question of how it may be possible to assess empirically the likely wealth effect on

consumption. Using a suitable panel technique we find that the hypothesis of the long-run mpc_w being the same across countries cannot be consistently rejected, and obtain a plausible estimate for the cross-section of eleven OECD countries. This estimate is a little over 6%, broadly consistent with estimates used in a wide range of policy models.

1 Introduction

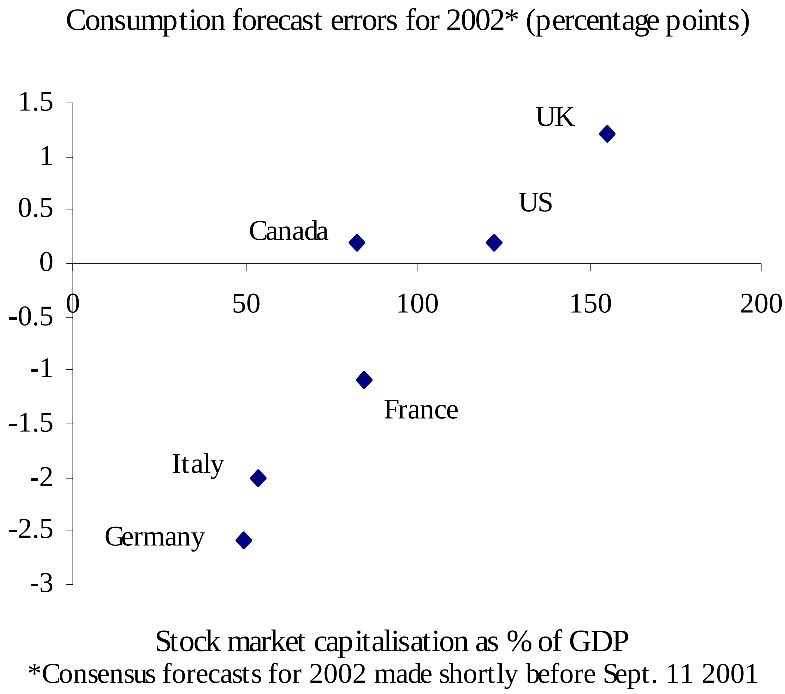
Over the past decade, there have been large changes in stock market capitalisation across the world. For example, between Autumn 1994 and Autumn 2000, stock market capitalisation increased by over 100% of GDP in the United Kingdom and the United States (see Chart 1). These rises reflected a trend in some countries towards privatisation, greater equity issuance and apparent upward revisions to expected future profits. Although slightly smaller in magnitude, similar increases occurred in the countries that are now part of the euro area. However, these large changes were partly reversed between 2000 Q3 and 2003 Q2, when capitalisation fell by over 40% of GDP in France, Germany and Italy, by 65% in the United States and by 87% in the United Kingdom.

Chart 1: Stock market capitalisation



Such changes in stock market capitalisation have brought wealth effects on consumption back onto the agenda. This reflects the fact that changes in stock market capitalisation can be a key channel through which changes in expected future income affects current consumption. Given that the post-2000 Q3 fall in market capitalisation in the United Kingdom and the United States was far higher as a percentage of consumption and GDP than in Germany and Italy, other things being equal one might have expected forecasts of UK and US consumption to have been revised down relatively more than for the other countries following declines in equity prices. Forecasts were indeed revised down more in countries where stock market capitalisation was relatively high, but outcomes were not in line with these forecasts. In January 2001, Consensus forecasts predicted that in 2002, consumption in Italy and Germany would each grow by more than 2.5%;[(1)] subsequently, forecasts were revised down to eventually be negative in each country. By contrast, after the initial downward adjustment, further revisions to US and UK consumption forecasts were small, notwithstanding the much more marked declines in market capitalisation. Chart 2 illustrates that forecast errors for consumption derived from Consensus projections tended to be larger during this period for those countries with lower stock market capitalisation.

Chart 2: Stock market capitalisation and Consensus forecast errors



What could be the reason for such differences in the pattern of forecast errors across countries? One explanation is that, in some circumstances, changes in stock market capitalisation may systematically understate wealth effects in lowly capitalised economies relative to those with high capitalisation. To the extent that the declines in equity prices during 2000 Q3 were driven by a downward revision to expected earnings of the corporate sector that was broadly symmetric across the major economies, the effect on consumption will ultimately be similar across countries. Yet the potential impact of the shock may be less immediately apparent in economies with low market capitalisation; the impact could come in large part through a change in the value of the unquoted

(1) See www.consensuseconomics.com.

sector.[2]

One lesson from this episode might be that partial equilibrium approaches to capturing the impact of changes in wealth on consumption face a cocktail of data problems and a failure to account for underlying structural causes of changes in both consumption and wealth. For example, in circumstances in which there are shocks to expected earnings, economies where market capitalisation is low and wealth held in unquoted equities is *under*-recorded, might be (inaccurately) estimated to have a *higher* mpc_w. Because conventional empirical estimates of the mpc_w are commonly calculated by dividing an empirical estimate of the partial elasticity of consumption with respect to wealth by the observed wealth to consumption ratio, if the reaction of consumption to an earnings shocks is similar, but wealth is under-recorded because of data problems, then the mpc_w will be overestimated. We provide some evidence for this effect in Section 4. What this means is that there are obstacles to estimating the effect of changes in market capitalisation (and financial wealth more generally) on consumption across countries: they illustrate the difficulties faced by forecasters and econometric models in capturing shocks that affect wealth and consumption.

The purpose of this paper is to offer a critique of the existing literature on wealth and consumption. Within this general objective, we focus on empirical estimates of the long-run marginal propensity to consume from financial wealth (mpc_w) and assess whether the magnitude of cross-country differences is plausible. We note at the outset that the long-run mpc_w is not necessarily a helpful concept. First, the question 'How much will consumption increase if wealth rises by £1?' is not well-formulated. There is a prior question about what caused wealth to rise in the first place; in general equilibrium, both wealth and consumption will respond to various shocks which may affect them jointly. As such, there is no such thing as an exogenous rise in wealth that subsequently causes an increase in consumption. Second, the consumer's budget constraint ensures that consumption will also have a mechanical relationship with wealth and, as such, we might not expect big variations across countries in the long-run response of consumption to changes in wealth broadly defined. This is not to say that the short-run mpc_w does not vary across countries - for instance because of liquidity constraints - but the extent to which such factors should have a marked influence over consumption in the long-run is questionable. As such,

(2) Which, as we discuss below, is also not measured consistently across economies. A further possibility is that there were additional coincident factors influencing consumption, such as country-specific changes in expectations about future labour income or changes in housing markets.

differences in the long-run mpc_w may reflect other influences, including cross-country data inconsistencies.

The structure of the paper is as follows. We regard the potential for cross-country data inconsistencies to be of such importance that it merits an up-front billing, so in Section 2 we describe potential wealth measurement issues in some detail, and present the wealth data used in the other studies and in this one. Section 3 presents the puzzle the paper seeks to address: existing empirical approaches deliver a wide range of estimates of the mpc_w across countries, whereas calibrated mpc_w are more similar across countries. Our own set of single-country SVAR-based estimates of wealth effects across countries tend to confirm that the mpc_w (apparently) differs widely across countries. In Section 4 we consider the possible explanations for those differences. In Section 5, we provide new evidence for the eleven countries using a panel technique that provides the means to explicitly test the hypothesis that the long-run mpc_w is the same across countries. Section 6 concludes.

2 How to measure wealth

Data problems have the potential to undermine the precision of all econometric estimates, but the potential mismeasurement of wealth across countries is particularly acute and the implications for model estimation and simulation properties are correspondingly severe. For this reason, we discuss data issues earlier and with greater prominence than is common in other research papers.

If the reaction of consumption to an earnings shock is similar, but wealth is under-recorded because of data problems, then the mpc_w will be overestimated. We show evidence for this effect in Section 4. Second, any simulation of wealth effects depends upon accurate measurement of the wealth-consumption ratio and any mismeasurement of this ratio can lead to difficulties, even in models that are calibrated rather than estimated. For example, the value of unquoted companies is not directly observable and, as we discuss below, there are significant cross-country discrepancies in statistical approaches used to obtain them. These statistical discrepancies will feed directly into inconsistencies in the measurement of wealth, and therefore into the simulation properties of any macro model.

In order to illustrate some of the anomalies in cross-country wealth data, Table A shows measures of wealth, broken down by asset type for a number of the major economies. Because of their sparsity, data at this level of disaggregation are not used for empirical estimates in any international comparison study of wealth effects (including this one), but they do suggest some potential anomalies in comparing data across countries. For example, the Belgian ratio of unquoted equity wealth to consumption is recorded as being higher than the German ratio of all directly held equity wealth to consumption.

Table A: Financial assets held by the household sector as percentage of consumption

Financial assets held by the household sector as percentage of consumption

	Net	**Gross, of which**								**Liabs.**
			directly held equities, o/w	quoted	unquoted	Mutual funds	Pensions & life ins.	Currency & deps.	Other	
US		**525**	196	126	70	49	150	64	66	
Japan		**493**	42	39	3	0	139	264	48	
Belgium	**456**	**539**	121	44	77	80	72	140	125	**83**
Netherlands	**426**	**596**	129				321	112	35	**170**
UK	**386**	**502**	90	52	38	26	265	100	21	**116**
Italy	**336**	**386**	107			65	47	93	73	**50**
France	**330**	**414**	154			36	90	110	23	**84**
Finland	**236**	**302**	189			7				**66**
Spain	**220**	**320**	108			40	41	116	16	**100**
Sweden	**209**	**311**	97			40				**102**
Portugal	**185**	**312**	73			26	43	145	25	**127**
Germany	**178**	**304**	48	37	11	34	84	104	35	**126**
Austria	**166**	**236**	40			9				**70**

All data are for 2000 except Belgium (2001). Data are from national sources and authors' calculations: data for financial liabilities are from Eurostat except Japan and the United States.

Differences across countries in the measurement of wealth can occur for a number of reasons. Babeau and Sbano (2002), for example, suggest three possible sources of inconsistency: first, the concept of wealth is not exactly the same across countries; second, there are errors in the measurement of wealth in some countries; and third, even where there are general guidelines on measuring wealth, the information necessary to apply them is not available in all countries. The next part of this section considers these three sources of inconsistency in turn, focusing on the measurement and application of guidelines on measuring wealth. The remainder of the section discusses the data used in other studies and in this one.

2.1 Some conceptual and data issues

2.1.1 Conceptualising wealth

At a fundamental level, an asset may be measured as financial wealth in one country but not in another. For example, sole proprietorships and partnerships are included as non-financial corporations in the US financial accounts whereas in Europe they are included in the household sector. This means that the net worth of sole proprietorships is regarded as households' financial assets in the United States, while in European countries part of the amount (professional buildings, plant and machinery) is regarded as non-financial assets. This will tend to increase estimates of net financial wealth in the United States relative to those for European countries.

2.1.2 Measurement of wealth

A second source of inconsistency across countries stems from problems in measuring wealth; for example there could be measurement errors in estimating the value of unlisted equities. Such problems are likely to be significant. In general, national accounts provide reliable data on households' holdings of shares listed on organised exchange markets. Holdings data are available from, for instance, centralised securities depositories or investment fund managers. The value of the holdings at any point in time can be easily calculated with reference to the price of the equity on the organised exchange. In the case of unlisted equities, however, there are difficulties with both: it is difficult to quantify households' holdings and, as we discuss below, it is difficult to value them. Yet unlisted equities are an important component of wealth, particularly in some euro-area countries: Babeau and Sbano (2002) note that at the end of 2000, the percentage of financial wealth held as unlisted shares was 18.7% in Italy, 19.8% in France and 22.7% in Spain, which is high in comparison with the Netherlands (5.3%), United Kingdom (8.5%) and Germany (9.7%). Norman, Sebastia-Barriel and Weeken (2002) show that unquoted equity accounted for almost half of equity holdings in France at the end of 2000.

2.1.3 Application of guidelines on measuring wealth

It is possible that the discrepancies in estimates of wealth across countries may be the result of differences in applying guidelines on measuring unlisted equities, and in the treatment of different unquoted companies based on size. In general, the valuation of equities follows the principles of

European System of Accounts (ESA95): marketable financial instruments are valued according to their market value. In the case of unlisted shares, however, it is not possible to calculate the market value with reference to the price at which the equity is traded on an organised exchange; holdings will typically be illiquid and traded infrequently. The general principle of ESA95 is to estimate the value of unlisted shares with reference to the value of listed shares in that country, taking into account the sector and the differences between the listed and unlisted companies due to different liquidity and reserves. However, the specific methodology applied by each country depends on the availability of these data.

Given the marked changes in stock market capitalisation in the late 1990s, the estimates of the value of unlisted equities at the end of 2000 in Babeau and Sbano (2002) may tend to overstate the importance of unlisted equities for some countries and understate it for others depending on the extent to which market prices are incorporated in the methodology. France and Spain apply market price criteria, ie criteria more closely related to stock market valuations. By contrast, in Germany unlisted public companies are evaluated with reference to the market value of listed companies but a heavy discount is applied and unlisted private companies are evaluated according to their book value. Therefore, there is a possibility that equity in unlisted companies is overestimated in France and Spain, and underestimated in Germany. Overall, we have little confidence that statistical treatments of wealth data are sufficiently harmonised to permit empirical estimation on a consistent basis; cross-country anomalies in statistical treatment of wealth form a major hurdle in comparing wealth effects.

2.2 Data used in this and other studies

This section discusses the data used in the empirical part of this paper and in other empirical studies; we focus on wealth data, but also discuss consumption and income data.

For wealth, in this paper we use net financial wealth data covering eleven countries. In most cases, data cover the period 1970 Q1 to 2002 Q4 though in some cases the sample is shorter. The data are originally sourced from OECD National Accounts, national statistical institutes and central banks, obtained via Datastream and the database of NiGEM, the global model developed by the National Institute of Economic and Social Research (NIESR). Net financial wealth is calculated as the sum of public debt, net foreign assets and miscellaneous assets including equities held by the personal sector, minus personal sector liabilities. In the case of net financial wealth data that are annual, we use the quarterly series interpolated by NIESR using quarterly information of the components of financial wealth where available. This measure is lagged one period to produce a measure of beginning-of-period wealth.

Other studies have used a wide range of data for financial wealth. Ludwig and Slok (2004) use stock market capitalisation data, the main advantage of which is that these data are readily available on a timely basis for a wide number of countries. Furthermore, capitalisation is measured accurately and changes in equity prices may capture changes in expected future income more rapidly than other forms of financial wealth. Yet the disadvantage of stock market capitalisation data is that they do not accurately capture domestic households' net financial wealth: in particular, domestically capitalised equities may be held by overseas households and by firms. Furthermore, household wealth is held in assets other than equities and the price of some of these assets may sometimes be expected to move in the opposite direction to equity prices (for example, equity prices tend to rise following a positive real demand shock, but bond prices tend to decline). Turning to other recent studies, Boone, Giorno and Richardson (1998) use equity holdings and Bertaut (2002) uses gross financial wealth based on national accounts data. Byrne and Davis (2003) use a compilation of wealth data that is close to the one used in this paper.

For consumption, we use total private consumption. The theoretically preferred measure would be consumption excluding durables as this measures the flow of services that is the basis of the utility to the consumer (see Deaton (1992) for discussion) though practical considerations complicate matters: Palumbo, Rudd and Whelan (2002) suggest that relating real consumption of non-durables to measures of real income and wealth, deflated by a price index of total consumption expenditures, is only valid if real consumption of non-durables is a constant measure of total real consumption. In any case, consumption excluding durables are not available for most countries, so we follow the practice in the majority of international studies by using a measure of total consumption.

For income, we use personal disposable income, which is calculated as labour income plus transfers net of taxes and interest, profit and dividend income. This measure of income implies double counting of income from wealth due to the inclusion of interest, profits and dividends; ie the income stream from financial assets, the present value of which defines wealth. The theoretically preferable measure of income would exclude interest profits and dividends in order to avoid double-counting income from financial wealth, but data difficulties prevent us from doing this for all countries across the full sample period and we therefore use personal disposable income as a proxy for this theoretically preferable measure. However, we have also explored two alternative income measures, the first of which excludes a residual measure of income, profits and dividends, and the second of which is labour income. We report results for each of these alternative measures in Section 3.4 and Section 5.

As the budget constraints are formulated in terms of real consumption, assets, income, and asset returns, each nominal variable in the budget constraint is deflated by the implicit household consumption deflator; because a common price index is used, the ratios of real variables are identical to the ratios of the corresponding nominal variables. All series are seasonally adjusted.

3 How different are mpc_w across countries?

3.1 Theoretical foundations

Most recent work on the wealth effect on consumption is based on the permanent income hypothesis (PIH), see for example Deaton (1992). In this model, households perceive an unanticipated, non-temporary increase in financial wealth as an increase in permanent income; which they use evenly over the remainder of their lifetime, boosting consumption in the current and all future periods. The PIH therefore implies that 'consumption is the annuity value of current financial and human wealth' (see Deaton (1992, page 80)).

In order to derive a measure of the wealth effect on consumption and formally derive the mpc_w, consider the conventional budget constraint for the consumer, which can be written as[3]

$$W_{t+1} = (R_{w,t+1})[W_t + Y_t - C_t] \qquad \textbf{(1)}$$

We use standard notation so that C_t and Y_t refer to consumption and 'labour' income in period t, and W_t refers to the stock of wealth at the beginning of period t; $(R_{w,t+1} = 1 + r_{w,t+1})$ is the return on the whole portfolio of assets. In words, this expression states that wealth in period $t + 1$ equals the resources available at period t (wealth plus saving in the current period) invested at the interest rate r.

(3) This section follows Altissimo, Georgiou, Sastre, Valderrama, Sterne, Stocker, Weth, Whelan and Willman (2005).

Solving **(1)** forward and imposing the terminal condition that at the end of the finite horizon the limit of discounted future wealth is zero, we obtain the intertemporal budget constraint

$$E_t \sum_{i=1}^{T} \frac{C_{t+i}}{\prod_{j=0}^{i} R_{w,t}^{-1} R_{w,t+j}} = W_t + E_t \sum_{i=1}^{T} \frac{Y_{t+i}}{\prod_{j=0}^{i} R_{w,t}^{-1} R_{w,t+j}} \tag{2}$$

According to this expression, the discounted value of planned future consumption is equal to today's total wealth, which is the sum of real and financial assets and the discounted sum of expected future labour income. This expression also implies that, as a response to a permanent wealth shock, the sum of consumption must rise by an equal amount, where Δ indicates the difference between the post-shock and pre-shock values. Assuming returns are constant over time, we are able to rewrite:

$$\Delta W_t = E_t \sum_{i=1}^{T} \frac{\Delta C_{t+i}}{\prod_{j=0}^{i} R_{w,t}^{-1} R_{w,t+j}} = \frac{1-\overline{R}_w^{-T}}{1-\overline{R}_w^{-1}} \overline{\Delta C} \tag{3}$$

This can be rearranged to give the mpc_w, ie

$$\frac{\overline{\Delta C}}{\Delta W_t} = \frac{1-\overline{R}_w^{-1}}{1-\overline{R}_w^{-T}} \tag{4}$$

where $\Delta \overline{C}_t$ refers to the average level shift in consumption over the time horizon from period t to T, and $\overline{R}_w$ refers to the average return on wealth between periods t and T. This expression highlights the two key determinants of the mpc_w: the time horizon over which consumers plan to spend the increase in wealth, and the return on financial assets.[(4)] [(5)]

3.2 The mpc_w in calibrated models

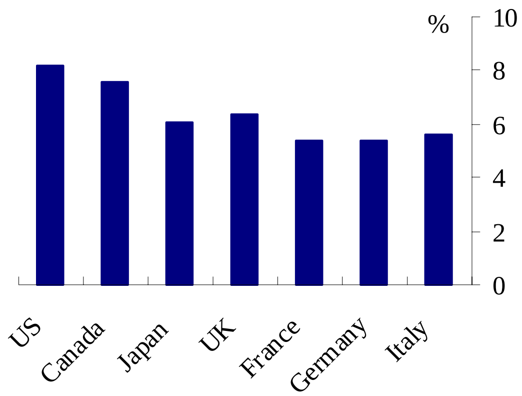
In calibrated models, such as the IMF's MULTIMOD, the euro-area dynamic general equilibrium model (EDGE) developed at the Finnish central bank, and the Bank of England Quarterly Model (BEQM), parameters such as the rates of return and the planning horizon are some of the deep structural parameters which are calibrated to the economies which they model. The mpc_w implied by these models therefore provide a theory-consistent guide to reasonable values that one might expect for the mpc_w.

As an illustration, Chart 3 shows the mpc_w in MULTIMOD.[(6)] These mpc_w estimates are based on

(4) There are other relevant factors once the coverage of wealth is extended to housing and there is habit persistence in consumption.
(5) When we examine theoretically plausible ranges for the mpc_w later in this paper, we report values of the mpc_w for a range of planning horizons and expected returns.
(6) For MULTIMOD, see Faruquee, Isard, Laxton, Prasad and Turtelboom (1998).

the Mark III version, as this provides separate mpc_w for the countries of the euro area. These mpc_w range between 5.4% and 8.2%. The highest values are those for Canada and the United States, followed closely by those for the United Kingdom and Japan; the values for the euro-area economies are somewhat lower. Overall, the mpc_w tends to be very similar across countries. In fact, three of the four parameters determining the mpc_w in MULTIMOD - the intertemporal elasticity of substitution in consumption, the real interest rate, and the probability of death - are the same across countries in the long run, as a result of calibration or pooled estimation (see Faruquee *et al* (1998)). The only parameter allowed to vary across countries, and hence to affect cross-country differences in the mpc_w, is taxes.

Chart 3: The mpc_w in MULTIMOD Mark III



3.3 The mpc_w in estimated models

A simple way to empirically estimate the mpc_w is to use a reduced-form consumption equation such as

$$C_t = \beta_0 + \beta_W W_t + \beta_Y Y_t + \varepsilon_t \qquad \textbf{(5)}$$

where C_t and Y_t are private sector consumption and income in period t, and W_t is wealth (measured at the beginning of period t). The coefficient on wealth in this equation provides an estimate of the mpc_w since $\beta_W = \frac{\Delta C}{\Delta W} = mpc_w$. In practice, however, estimation of equation **(5)** may not be feasible if, as occasionally found in practice, the right-hand side variables prove to be non-stationary and do not cointegrate with the regressand. In this case, estimating the equation may produce a spurious regression.

In the annex, we present evidence on the series used in the empirical part of this paper. For these series, the evidence suggests that individually, consumption, wealth and income appear to be non-stationary, but the panel evidence points to stationarity, which means that for the data we use in this paper estimating equation **(5)** may not lead to any problems. In the annex, we also present evidence on panel co-integration which appears to support this proposition.

In the literature, two other types of consumption equation have been used, each based on a particular transformation of the variables in **(5)** and the estimation of an equation specified in terms of the transformed variables. These types of equations have sometimes been presented as providing an appropriate transformation to ensure stationarity, although this may not always be the case. Moreover, the alternative transformations suggested in the literature affect how the estimate of the mpc_w is obtained. In the following, we discuss the implications of using other types of consumption equation to estimate the mpc_w.

3.3.1 Estimates of the mpc_w from a ratio specification

The first of the transformations used in the literature consists in dividing equation **(5)** by income. Assuming that β_0 is zero, this yields a regression of the form

$$\frac{C_t}{Y_t} = \beta_Y + \beta_W \frac{W_t}{Y_t} + \varepsilon_t \qquad \textbf{(6)}$$

which relates the consumption-income ratio to the wealth-income ratio.(7) Palumbo *et al* (2002) show that (for the case of a constant interest rate) the PIH implies that there should be a stable long-run relationship between consumption-income and wealth-income ratios, so that this equation can be estimated without problems.

In the literature, equation **(6)** has been estimated on alternative series for each of the variables it includes. Based on this specification, the literature has generally found that coefficients are higher for consumption excluding durables than total consumption, and higher for disposable income compared with labour income. With respect to different wealth measures, the literature has found higher coefficients for equity than non-equity wealth. The latter result has also been found based on estimation of an equation providing a joint estimates of the mpc_w for different wealth

(7) Note that the coefficient in this regression is only approximately equal to the mpc_W, as this regression uses an approximation, as the estimated coefficient is $\beta_W = \frac{C_t - C_{t-1}\frac{Y_t}{Y_{t-1}}}{W_t - W_{t-1}\frac{Y_t}{Y_{t-1}}} \neq \frac{\Delta C}{\Delta W_t} = mpc_w$.

measures, as in

$$\frac{C_t}{Y_t} = \beta_Y + \sum \beta_{W_i} \frac{W_{i,t}}{Y_t} + \varepsilon_t \qquad \textbf{(7)}$$

where i is the index for the components of wealth (eg financial and non-financial (physical) wealth, or equity and non-equity wealth).[8]

3.3.2 Estimates of the mpc$_w$ from a log-linear specification

Most of the existing empirical work on cross-country comparisons though has used a logarithmic transformation, resulting in an estimable equation of the type

$$c_t = \beta_y y_t + \tilde{\beta}_w w_t + \varepsilon_t \qquad \textbf{(8)}$$

where lower-case letters denote logarithms of variables. Due to the logarithmic transformation, this regression provides an estimate of the elasticity of consumption with respect to wealth, $\tilde{\beta}_w = \frac{\Delta C_t / C_t}{\Delta W_t / W_t}$, of which the mpc$_w$ is a transformation. In order to obtain an estimate of the mpc$_w$ the elasticity has to be multiplied by the consumption-wealth ratio, such that

$$\tilde{\beta}_w \frac{C_t}{W_t} = \frac{\Delta C_t / C_t}{\Delta W_t / W_t} \frac{C_t}{W_t} = \frac{\Delta C_t}{\Delta W_t} = mpc_w \qquad \textbf{(9)}$$

Equations **(8)** and **(9)** illustrate that the long-run mpc$_w$ may vary across countries owing to differences across countries in the elasticity of consumption with respect to wealth and in the wealth-consumption ratio.

There are several potential shortcomings of estimating a log-level specification. The first becomes important if it is assumed that the mpc$_w$ differs across the various types of wealth and across different components of consumption (eg durables and non-durables). This can be easily shown by expressing consumption as a basket of goods (here indexed by j), $C = \sum_j C_j$, and wealth as a portfolio of assets (here indexed by k), $W = \sum_k W_k$, such that

$$\mu_{C,W} = \frac{\sum_j \Delta C_j}{\sum_k \Delta W_k} = \sum_j \alpha_k \mu_{C_j, W_k} \qquad \textbf{(10)}$$

and

$$\eta_{C,W} = \frac{\sum_j \Delta C_j}{\sum_k \Delta A_k} \frac{\sum_k W_k}{\sum_j C_j} = \sum_j \gamma_j \sum_k \alpha_k \eta_{C_j, W_k} \qquad \textbf{(11)}$$

where

$$\gamma_j = \frac{C_j}{\sum_j C_j} \text{ and } \alpha_k = \frac{W_k}{\sum_k W_k} \qquad \textbf{(12)}$$

(8) However, one potential difficulty of estimating this type of equation is that the estimate of the mpc$_w$ may reflect differences in the income elasticities of the demand for different assets.

and where $\mu_{C,W}$ and $\eta_{C,W}$ denote the mpc_w and the elasticity of consumption with respect to wealth, respectively.

Equation **(11)** shows that the elasticity estimate compounds the wealth effects of different assets and parts of the consumption basket. So while estimating the effect of aggregate wealth on aggregate consumption should still provide a useful benchmark, ideally empirical work should attempt to estimate wealth effects on consumption for different assets and goods, especially in an international context, since consumption baskets (γ_j) and portfolios shares (α_k) are likely to differ across countries and because both assets and goods differ in their characteristics, so that consumers may not respond in the same way to changes in every component of wealth and not with all parts of their consumption. Assets, for example, differ in rates of return and volatilities, as well as the extent to which they can serve as collateral to credit-constrained consumers, and goods differ in the time horizon over which they provide utility.

Another limitation of the logarithmic transformation has recently been illustrated by Rudd and Whelan (2002) in their comment on an approach followed by several authors, eg Lettau and Ludvigson (2001). This approach uses a log-linear specification starting from the equation

$$c_t - w_t \approx E_t \sum_{i=1}^{\infty} \rho_w^i (r_{w,t+i} - \Delta c_{t+i}) \qquad \textbf{(13)}$$

which is derived from an intertemporal budget constraint relating consumption to wealth, imposing a transversality condition on wealth and - to ensure it holds *ex ante* as well as *ex post* - taking expectations (for details of the derivation, see Rudd and Whelan (2002)). One key feature of this equation is that it subsumes labour income into aggregate wealth. By assuming that labour income approximates the non-stationary component of human wealth, and that the return on aggregate wealth is a weighted average of the returns to human and other wealth, this equation can be rewritten as

$$c_t - \omega a_t - (1-\omega) y_t \approx E_t \sum_{i=1}^{\infty} \rho_w^i (\omega r_{a,t+i} + (1-\omega) r_{h,t+i} - \Delta c_{t+i}) + (1-\omega) z_t \qquad \textbf{(14)}$$

where z_t is a mean-zero variable capturing the stationary part of the deviation between labour income and human wealth. If the RHS terms are stationary, as assumed in Lettau and Ludvigson (2004), then the LHS needs to be stationary too, so that a cointegration analysis of wealth, consumption and labour income can be used to assess to what extent fluctuations in these variables are caused by permanent or transitory shocks.[9] However, as pointed out by Rudd and Whelan

(9) As pointed out by Lettau and Ludvigson (2001), this has far-reaching implications for traditional macro models

(2002), equation **(14)** is strictly applicable only to aggregate consumption which, given our discussion above of differences in the mpc_w across types of assets and components of consumption, makes an analysis based on this equation much less useful.

3.3.3 Estimates of the mpc_w from the empirical literature based on log-linear specification

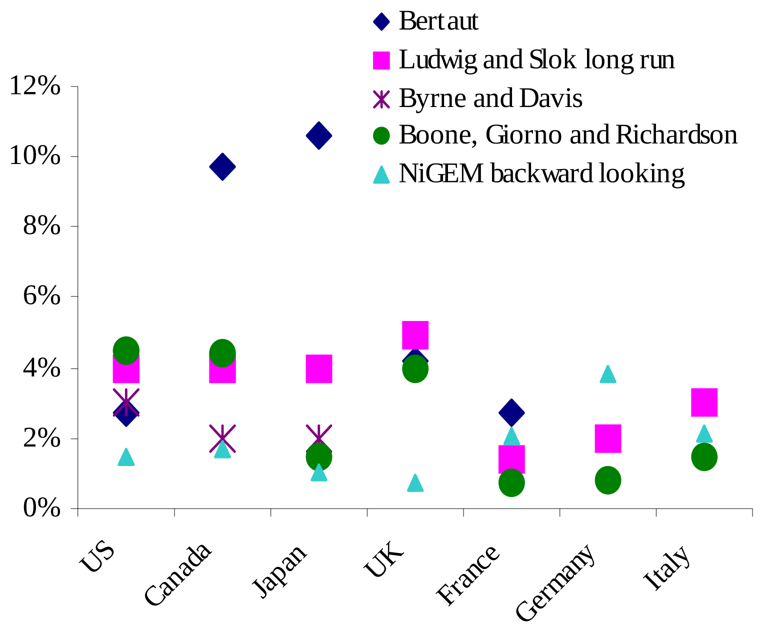
In Chart 4, we show existing estimates of the mpc_w based on log-linear specifications, obtained from three recent contributions to the literature for the major seven economies. The chart also shows mpc_w from a study in which they are calculated indirectly from an assumed estimate of the mpc_w for the United States and scaled for other countries based on the households' equity holdings to income. It shows a fifth set of estimates, derived from the backward-looking versions of consumption equations taken from the NiGEM model, an empirically estimated model whose reduced-form consumption equations are restricted to conform with the permanent income hypothesis (in the long-run coefficients on wealth and income are constrained to sum to one);[10] in order to present these estimates in the same form as those used in the four studies included in Chart 4, we translate the elasticities into mpc_w using the wealth-consumption ratios described above and used later in this paper. The mpc_w estimates presented in the chart are long-run; the speed of adjustment to the new equilibrium will vary across economies.

Chart 4 demonstrates a large variation in the empirical estimates, both comparing the same country across different studies and comparing across countries. How surprising are these variations? The studies included in Chart 4 differ in their approach to deriving the mpc_w. Bertaut (2002), Ludwig and Slok (2002), and Byrne and Davis (2003) use the same method of calculating the mpc_w via estimating the elasticity of consumption with respect to wealth from equations **(8)** and **(14)**, and then using the wealth-consumption ratio to give the mpc_w. Within this approach Ludwig and Slok (2002) estimate elasticities of consumption with respect to wealth for two groups of economies: so-called bank-based economies are identified as including France, Germany, Italy and Japan, while the group of so-called market-based economies is defined to include Canada, the United Kingdom and the United States. Boone *et al* (1998) do not estimate

assuming constant discount factors. Given that consumption and income are not very volatile in the short run (and unlikely to have higher volatility in the long run), the only way that these three variables can cointegrate is through low volatility of asset prices in the long run. This implies that shocks to asset prices must be transitory and that asset prices must be predictable, which a growing empirical literature tends to support. However, this can only be the case if markets are inefficient, or if discount rates are time-variable (which has been argued in the finance literature for some time).

(10) For NIGEM, see Barrell, Dury and Pain (2001).

Chart 4: Estimates from the literature of the long-run mpc$_w$



elasticities for each country, instead assuming an mpc$_w$ for the United States of 4.5 cents per dollar, which is then scaled for the other countries according to the country's ratio of households' equity holdings to income relative to the ratio in the United States. The Boone *et al* (1998) result is therefore driven by the fact that the US ratio of households' equity holdings to income is similar to that of Canada and the United Kingdom and considerably higher than the euro-area countries.

Across the studies cited above, there are also differences regarding the period over which the elasticity is estimated and the wealth-consumption ratio is calculated. Bertaut (2002) uses different sample periods for different countries, ranging from 1960 to 2000 for the United States to 1978 to 1998 for France, and the wealth-consumption ratio is calculated as an average over 1995 and 1998. Ludwig and Slok (2002) use a sample period of 1985-2000 to calculate elasticities and the most recent wealth-consumption ratio. Byrne and Davis (2003) use 1970-2000 as a sample period and calculate the average wealth-consumption ratio over this period. Boone *et al* (1998) use ratios of households' equity holdings to income in 1997. As we discuss below, the mpc$_w$ varies in line with the wealth-consumption ratio which is typically not constant across countries and time periods.

Such differences in methodology and data may explain some of the cross-country variation in mpc$_w$ across different studies. But, even within individual studies, there are marked differences in the estimated mpc_w across countries. Before assessing the plausibility of these cross-country differences, we therefore present new estimates of the mpc_w for a cross-section of eleven OECD countries.

3.4 New estimates of the mpc_w based on SVARs

Our cross-section of eleven OECD countries includes Belgium, Canada, France, Germany, Italy, Japan, the Netherlands, Portugal, Spain, the United Kingdom and the United States. The approach we take uses the estimates of the elasticity of total consumption with respect to net financial wealth obtained from structural VARs. Structural VARs have the advantage of explicitly allowing for feedback effects from consumption to wealth, something that the single-equation literature presented in the previous section cannot properly address.[(11)] Furthermore, the structural VAR approach has the potential to demonstrate how the response of consumption and wealth vary according to the nature of the shock that affected them.

We use a total of four different VARs. The first two VARs are adapted from Lettau, Ludvigson and Steindel (2002) (LLS in the charts below), where they have been used to analyse the importance of wealth effects in the transmission of monetary policy. The first has five variables, which are: consumer price inflation, labour income, consumption, total financial wealth and the interest rate. The second VAR also includes a general index of commodity prices[(12)] and thus has six variables in total. Both VARs are estimated in (log) levels.[(13)] As a check for robustness, we also estimate much simpler tri-variate VARs of consumption, income and total financial wealth, both in (log) levels and differences.

Following Lettau *et al* (2002), the five and six-variable VARs are identified by assuming that (i) inflation and income respond to other variables with a lag (and inflation responds to output with a lag), (ii) the stock of wealth at the start of a period does not respond to the flow of consumption during that period and (iii) the interest rate does not respond directly to wealth. The assumptions (i) and (ii) reflect common assumptions about a minimum delay in the reaction of macro variables

(11) The importance of taking a systems approach to estimating wealth effects has also been emphasised in a recent study by Blake, Fernandez-Corugedo and Price (2003).

(12) These data are downloaded directly from the Commodity Research Bureau, the same source as used in Lettau *et al* (2002).

(13) The estimation in levels yields consistent estimates also in the case of non-stationary variables if there are cointegration relationships between the variables.

to the policy instrument, and assumption (iii) reflects the fact that in practice asset prices only affect the policy instrument to the extent that they have an impact on inflation and output.[14] The three-variable VARs only require assumptions (i) and (ii), as they have no interest rate variable.

All four VARs have the form

$$B_0 z_t = B_1 z_{t-1} + .. + B_p z_{t-p} + u_t = \sum_{k=1}^{p} B_k z_{t-k} + u_t \qquad \textbf{(15)}$$

where the matrix B_0 reflects the simultaneous interactions between the variables in z_t, and any short-run restrictions on these interactions. The effects of a shock on the variables in z_t are given by the moving average (MA) representation relating the variables to the shocks,

$$z_t = A(L) u_t$$

where $A(L) = \left(B_0 - L^k \sum B_k z_{t-k}\right)^{-1}$. From this, it is possible to extract the elasticity of consumption with respect to wealth. In the case of the (log) level VARs, this is done by picking from the matrix $A(L)$ the elements which give the effects of a wealth shock on consumption and (for normalisation) the initial effect of this shock on wealth itself. The elasticity can then be converted into the mpc_w by using the wealth-consumption ratio. The procedure is similar for the difference VAR, once the responses for the log differences have been converted into the levels responses.

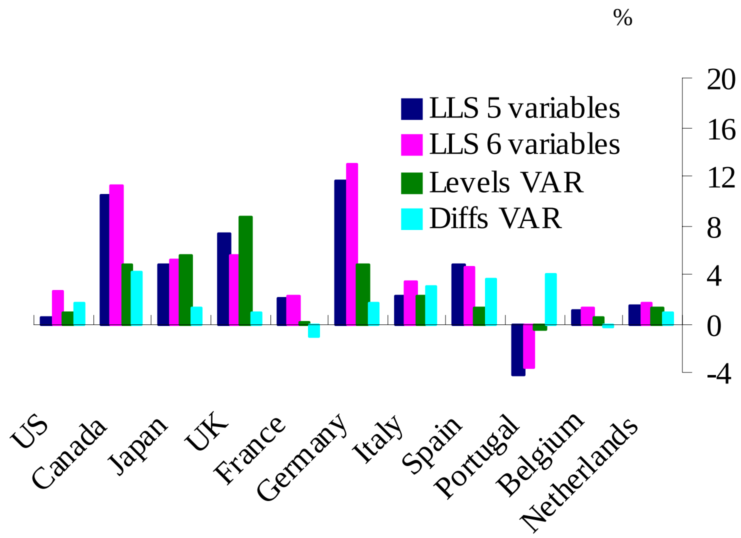
Chart 5 shows the estimates of the mpc_w, using the wealth-consumption ratio for 2001-02. They range between 1% and 5% for most of the euro-area countries in our data set. However, there are some surprising results: the mpc_w is markedly higher for Germany and Canada, and in some VARs there are even some negative estimates in the cases of the smaller countries and France. This is hard to rationalise on economic grounds and may in part reflect problems with data measurement.

More generally, there appears to be large cross-country dispersion in the empirical estimates of the mpc_w. And this conclusion appears to be sustained through a number of robustness checks. First, it appears to carry through irrespective of the choice of VAR. Second, it holds when we recompute the mpc_w using the wealth-consumption ratio for the full sample (1970-2002). These ratios tend to be smaller, although in the case of three countries (the Netherlands, Portugal and Spain) they are actually larger, either because wealth has not increased to the same extent or has not fallen by as much as in the case of the other countries. These results are shown in Chart 7. Such wide variation across countries is in line with existing literature that used single-equation estimates.

(14) See Lettau, Ludvigson and Steindel (2002) for the arguments supporting these restrictions.

Finally, we find a similar dispersion when we use the alternative income measures (disposable income adjusted for net taxes, as well as labour income, which excludes all income from assets, as well as labour income): the results were very similar to those presented above using the disposable income measure.

Chart 5: Empirical results, based on SVARs (using 2001-02 wealth-consumption ratios)

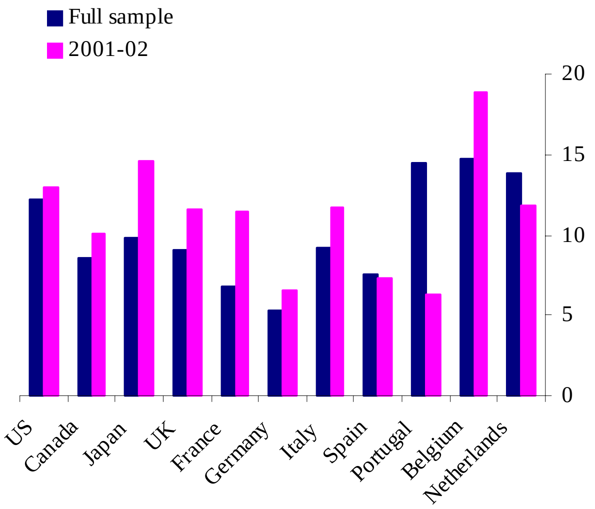


We end this section by emphasising that our new VAR-based estimates provide estimates of the mpc_w that appear to be sensible only when averaged across countries. As in the case of previous empirical estimates, the variation in our estimates appear to be large across countries, particularly when compared with those found in theoretically founded models. In the next section, we examine whether such a variation is plausible.

4 How plausible are differences in mpc_w across countries?

In this section we investigate a number of aspects of the plausibility of cross-country patterns in the mpc_w observed in the literature. For instance, we noted in the previous section that calibrated mpc_w are similar across countries, whereas empirical estimates of the mpc_w across countries and studies vary widely. In this section we seek to judge which is more accurate. Moreover, the literature has reported cross-country patterns in the mpc_w: some authors, for example, have found it to be higher in the United States than, say, Germany. We investigate whether this 'stylised fact' is well-grounded.

Chart 6: Wealth-consumption ratios

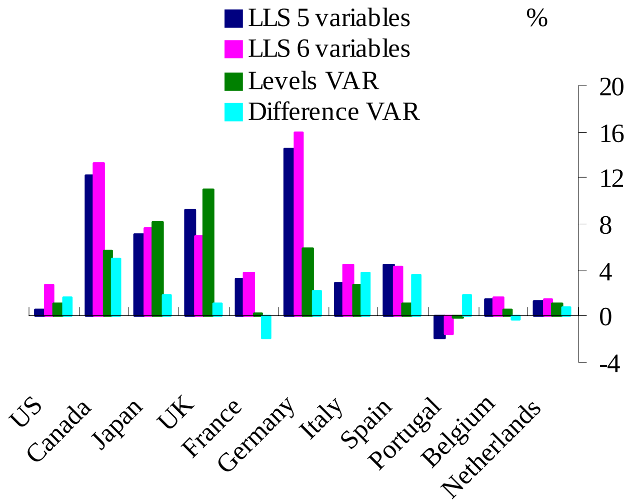


The approach we take is to identify and discuss a number of potential explanations of wide cross-country differences in empirically estimated mpc_w. First, it is possible that there are differences across countries in the length of the planning horizon or the rate of return on financial assets. Second there may be differences in the form of wealth holding. Third, we look at the role of housing. Fourth, the mpc_w is a reduced-form concept and there may be differences across countries in the source and duration of shocks. Finally, any approach to estimating mpc_w may be prone to errors arising from data mismeasurement.

4.1 The role of demographics

Differences across countries in the length of the planning horizon, reflecting deeply embedded structural and demographic differences, or the rate of return on financial assets, could in principle explain cross-country differences in the mpc_w. Below, we use cross-country data on wealth holdings in different income groups to calibrate a numerical example that shows such factors might explain only limited cross-country variation in the mpc_w. We also consider the extent to which cross-country variation in the mpc_w may be caused by differences in the concentration of ownership in particular types of asset, such as pension funds or bank deposits. We find limited support for the view that these are of sufficient magnitude to explain the size of the cross-country differences in the mpc_w.

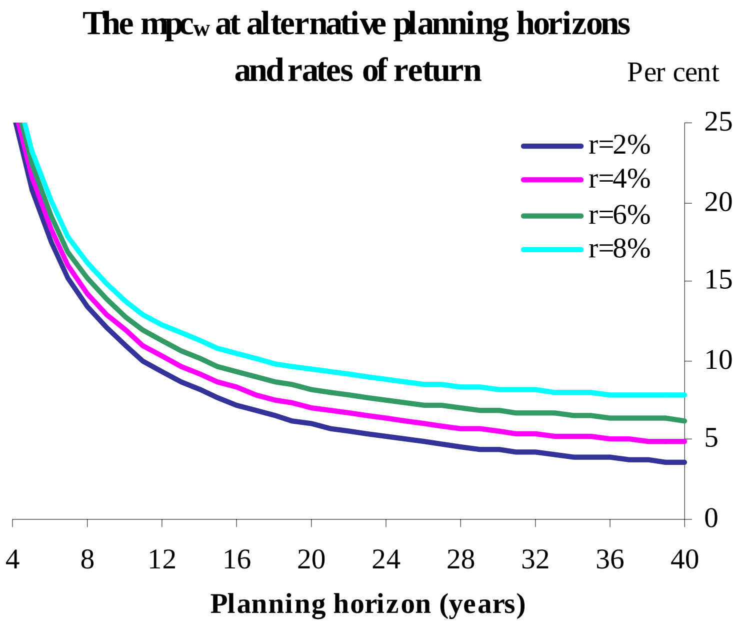
Chart 7: Empirical results, based on SVARs (using 1970-2002 wealth-consumption ratios)



Numerous authors, including Norman *et al* (2002) and Poterba and Samwick (1995) have focused on demographic factors as being of potential importance in explaining variation in the mpc_w across countries and time. In this section, we conduct a simple experiment based upon the formula for the mpc_w in equation **(4)**. We use data on the distribution of income and wealth, and a range of rates of return, to show how such differences across countries can affect mpc_w. The basis for the experiment is provided by Poterba (2000) who provides benchmark calculations of how a household may increase consumption following a favourable wealth shock. For a household with no bequest motive, the amount by which a household could increase consumption in the remaining years of its life depends on its planning horizon (life expectancy) and the real interest rate. Chart 8 presents benchmark calibrations of the amount by which a household might increase its consumption in response to an exogenous increase in wealth, for various expected returns and planning horizons. Each line shows, for a particular expected return, by how much the mpc_wvaries as the horizon over which an increase in wealth is consumed lengthens. The chart shows that as the planning horizons tend to infinity, the mpc_w tends to the real interest rate.

Based on this, we design an experiment in which the aim is to determine if demographic factors can lead to cross-country differences in the mpc_w through differences in planning horizons. To be able to reject the possibility that demographic explanations are important, we make the assumption that demographics strongly affect planning horizons. To be precise, we assume that low-income consumers consume a shock to their wealth over a much shorter planning horizon than do high-income consumers. This may occur because lower-income consumers are more likely to face cash-flow constraints.[15]

Chart 8: The mpc_w at alternative planning horizons and rates of return



We next illustrate how known cross-country differences in demographics might affect the mpc_w. The steps in this experiment are shown in Table B. In the first step, we assume that low-income households plan to spend any increase in wealth over much shorter time horizons than wealthier consumers. Row B illustrates our assumptions that low-income earners have a planning horizon of six years, whereas high-income earners plan to spend changes in their wealth over a 24-year horizon. Assuming that all consumers discount at a rate of 2% per annum, row C illustrates the implication that low-income consumers have an mpc_w of 18%, whereas high-income earners have an mpc_w of 5%.

In practice, are differences in the concentration of equity holdings across countries sufficient to justify a marked variation in mpc_w? In order to judge this we use data for demographic differences in the distribution of wealth across countries taken from Norman *et al* (2002). These data tend to show that, for instance, Italian holdings of directly held equity are more concentrated among high-income earners than in the United Kingdom. But do such differences have a marked impact on the mpc_w? In order to arrive at the results we are forced to make simplifying assumptions,

(15) There are other demographic factors - such as age profile - that might also be different across countries, but these do not exhibit such variance across countries as is the case for the distribution for equity wealth and income, so are less likely to generate significant dispersion across countries.

reflecting the paucity of comparable cross-country data on the distribution of wealth within economies.[16] Based on their data, step 2 of Table B illustrates our calculations of the extent to which directly held equities are more concentrated in the hands of high-income earners in Italy than in France, the United Kingdom and the United States. Under plausible assumptions, the data suggest that 79% of Italian directly held equity wealth is concentrated in the highest income quartile in Italy, whereas in the United Kingdom the same group hold only 55% of directly held equity wealth.

Table B: Distribution of wealth and cross-country differences in the mpc_w

Step 1: Calculate mpcW for various planning horizons					
A. Quartiles in income distribution	1-25	26-50	51-75	76-100	
B. Assumed planning horizon (years)	6	12	18	24	
C. Implied mpc_w with 2% real int. rate	0.18	0.09	0.07	0.05	
Step 2: Use wealth distributions in each country to estimate overall mpc_w					
Estimated proportion of direct equity wealth held by each income quartile and implied mpc_w					
Quartiles in distribution	1-25	26-50	51-75	76-100	**Implied mpc_w**
France	0.05	0.13	0.24	0.58	**6.6%**
Italy	0.01	0.04	0.16	0.79	**5.7%**
UK	0.05	0.11	0.29	0.55	**6.6%**
US	0.02	0.11	0.29	0.58	**6.3%**

Source: Norman *et al* (2002) and authors' calculations.

However, we do not find that these differences in the distribution of wealth lead to marked cross-country differences in the aggregate mpc_w. Even under the assumptions that planning horizons vary starkly across income quartiles, the aggregate mpc_w is only 0.9 percentage points higher in the United Kingdom than it is in Italy (the countries which, under the various assumption in this example, have the highest and lowest calibrated mpc_w). The similarities across countries in the mpc_w calibrated above are in stark contrast to the very marked divergences in the estimated mpc_w in the empirical literature (shown in Chart 4) and in our own estimates based on SVARs (shown in Chart 5). Furthermore, it appears from these data that the distribution of wealth among income cohorts would imply a marginally higher mpc_w in France than in the United States, in contrast to most theoretical calibrations (for example, those in MULTIMOD and in empirical estimates). The main lesson we draw from this analysis is that differences in the distribution of

(16) Norman *et al* (2002) provide data for the proportion of households holding wealth in various percentiles in income distributions that are uneven within and across countries. In order to derive the distribution of wealth according to income quartiles we first assume that wealth is held uniformly within all percentiles in income groups they provide. This enables us to estimate the proportion of households owning equities each income quartile. We then assume that low-income earners' holdings of equity are 25% that of high-income earners; low-mid earners 50%, and mid-high income earners 75%. This enables us to deduce the proportion of directly held equity wealth held according to income quartiles as shown in Table B.

equity wealth do not appear to justify marked cross-country differences in estimates of the mpc_w; moreover, the countries commonly estimated or calibrated to have higher mpc_w are not necessarily those that accord with our estimates from this experiment.

4.2 The form of wealth holding

Table A demonstrated marked cross-country differences in the type of financial wealth held. Given this, were the mpc_w to be higher on some assets than others, this might provide another justification for large cross-country differences in the mpc_w. A theoretical foundation for this argument is provided by Thaler (1994) and Thaler (1990). In his model, households divide wealth into a current income account (with a high mpc_w), a future income account (mpc_w close to zero) and an asset account (mpc_w in between the other two). To the extent that equity wealth is allocated to the asset account, this might explain the large response of consumption; it is also consistent with empirical evidence showing capital gains on equities tend to be saved, whereas one-off cash windfalls from equities (for example, from takeovers) tend to boost consumption.

However, the empirical case for differing wealth effects according to the type of asset is unproven. Poterba and Samwick (1995) for the United States considers whether the increasing proportion of indirect ownership of equities through pension funds has affected the link between stock market returns and consumption. In both cases there is little evidence that direct ownership of equity wealth has a different impact from indirect ownership on consumption. This may appear surprising given the relative visibility and liquidity of directly held equities relative to other forms of equity. However, it is possible to argue that there are no significant cross-asset differences in mpc_w. Consumers might be expected to have a lower mpc_w for assets whose value is volatile, such as equities, or illiquid, such as pension funds and unquoted equities. But as a direct result of such uncertainty and illiquidity, these assets typically yield higher rates of return (for example because of the equity risk premium) which in turn implies a higher mpc_w. For example, pension funds may be both illiquid and their value volatile. Conversely, they have historically offered high, tax-exempt rates of returns and it is unclear as to whether the offsetting effects should lead them to have a higher or lower mpc_w relative to other assets.

4.3 A role for housing?

In this section we assess the potential role for housing wealth. Various studies have found this to be important (see Case, Quigley and Shiller (2001), Dvornak and Kohler (2003) and Ludwig and Slok (2004)). But the effects are not the same across studies. And within studies, the effects vary over time. We shall argue that differences across countries might imply variations in the response of consumption to shocks but that the theoretical grounds for believing there is a long-run housing wealth effect across countries are uncertain, and the data deficiencies undermine efforts to detect any such effect.

In the case of housing, and in contrast to financial wealth, it is possible to identify differences across countries which suggest there could be systematic cross-country divergences in the response of consumption to shocks. To the extent that shocks to housing affect consumption in the short term through their impact on mortgage equity withdrawal, it is likely that such effects would occur more in countries in which the retail financial system allows comparatively cheap and easy re-mortgaging and/or taxes on housing turnover are low. Table C, where the data are taken from MacLennan, Muellbauer and Stephens (1998) and may be out of date for some countries,[17] illustrates stark differences in housing taxation in the EU. In Belgium, Greece and Portugal, stamp duty is 10% or higher so the incentive to move and withdraw equity from housing is much reduced. These reasons illustrate why the short-run mpc from housing may differ across countries.

Table C: Differences in housing taxation in EU economies

	Tax on capital gains	Stamp duty (%)	VAT on new homes (%)
Austria	If resold in a fixed period	3	10 to 20
Belgium	Yes	12.5	6 to 20.5
Denmark	If resold in a fixed period	1	25
Finland	If resold in a fixed period	6	22
France	If resold in a fixed period	7	18.6
Germany	If resold in a fixed period	2	0
Greece	No	10	n/a
Ireland	Rollover relief applies	up to 9	12.5
Italy	No	4.2 - 8	4
Netherlands	Yes	6	17.5
Portugal	Rollover relief applies	10	5 to 16
Spain	Rollover relief applies	6	7
United Kingdom	No	1	0

Source: MacLennan, Muellbauer and Stephens (1998).

More generally, in the short run there are a number of reasons why changes in house prices may

(17) Since 1997, higher rates of stamp duty have been introduced in the UK for transactions exceeding £250,000 and £500,000.

lead changes in consumption. For instance, a positive shock to permanent income would be expected to boost consumption of both housing and non-housing goods, but to the extent that house prices are more flexible than the price of other goods, this could imply house prices are observed to Granger cause consumption. There may also be credit market effects: houses represent collateral for home-owners (see Aoki, Proudman and Vlieghe (2002)), and a rise in house prices provides more collateral against which households can borrow, for instance in the form of mortgage equity withdrawal.

However, in the very long run it is less clear that changes in the aggregate value of housing wealth should lead to changes in aggregate consumption. Aoki *et al* (2002) argue that there is no such wealth effect on consumption. Households tend to live in their home and consume the housing services it provides, so the positive effect of higher house prices is offset by an increase in the opportunity cost of these housing services (an effect that is unlikely to be captured adequately in the consumption deflator). Furthermore, while changes in house prices redistribute wealth between agents in the economy, it is not clear that any negative impact of house price falls on the expected permanent income of households owning their house outright will always outweigh its positive impact on the permanent income of those servicing a mortagage (see King (1994)).

In any case, a practical obstacle to assessing cross-country differences in the impact of changes in housing wealth on consumption is data quality. Ludwig and Slok (2002) use house price data compiled by the Bank for International Settlements (BIS). These data represent the price rather than the value of the housing stock. Bertaut (2002) uses data for non-financial wealth for the relatively few countries where they are adequately available. However, like financial wealth data, non-financial wealth data are subject to cross-country differences in statistical treatment: for instance, in some countries national indices are available whereas in others there are only data on house prices in certain cities and metropolitan areas.

In summary, although differential taxation treatment across countries might imply systematic cross-country differences in the role of the housing market in the transmission mechanism, the theoretical grounds for believing there is a long-run housing wealth effect across countries are uncertain, and the data deficiencies undermine efforts to detect any such effect. We have explored specifications including housing wealth but have found these to be unsatisfactory. The results had different signs across countries and did not improve diagnostic results.[18] In sum, we have found data difficulties with regard to housing insurmountable and we do not find it surprising that previous literature differs in its views as to the relative role of housing effects across countries and time. We do acknowledge, however, that although housing is difficult to incorporate in econometric analysis, it may be an important omitted variable, at least in the short run (see footnote 1).

4.4 The duration and source of shock

A further source of potential difficulties in comparing wealth effects across countries arises because cross-country comparisons of wealth effects so far considered in this paper are reduced form, partial equilibrium. Such equations can inform us as to how, on average, consumption has responded to changes in equities (and financial wealth); but in general equilibrium both variables will respond to common shocks, which may affect wealth and consumption together.

The analysis of Lettau and Ludvigson (2004)[19] suggests that the perceived duration of shocks may be important in determining the responses of wealth and consumption. They argue that deviations in consumption from equilibrium may follow from time variation in the real interest rate and hence in equity prices. They show that large, transitory swings in equity prices have limited effects on consumption and find that the vast majority of the variability in consumption is driven by permanent shocks, whereas 88% of the variability in wealth is driven by transitory shocks. They find that transitory (albeit persistent) shocks to wealth that are driven primarily by volatility in equity prices do not significantly affect consumption. The authors produce alternative (lower) estimates of the mpc_w that take account of the fact that only permanent changes in wealth have a significant impact on consumption. Blake *et al* (2003) find qualitatively similar results for the United Kingdom. The remainder of this section assesses the likelihood that differences in the pervasiveness and the source of various structural shocks might affect estimates of the reduced-form relationship between consumption and wealth, and the extent to which we believe such considerations affect overall estimates of the mpc_w and cross-country differences in the estimates.

(18) Results available from the authors on request.
(19) Using techniques developed by Stock and Watson, they decompose the variables in a cointegrating relationship between US consumption, income and wealth into changes that are related to permanent or transitory shocks. Rudd and Whelan (2002) cast some doubts on the extent to which the intertemporal budget constraint should yield a cointegrating relationship once appropriate deflators are used for all variables.

Different shocks may imply a range of simultaneous reactions of consumption and wealth, yet numerous authors have demonstrated the intuitive result that consumption and wealth can respond in differing ways depending upon the nature of the shock. There are a number of other shocks that may lead to low correlations between wealth and consumption. Millard and Wells (2003) develop a simple general equilibrium model that suggests shocks to equity risk premia that result from increases in equity volatility will have no effect on consumption, whereas other shocks each induce a positive correlation between equity prices and consumption. Lantz and Sarte (2001) consider consumption and wealth in a general equilibrium setting and find that a change in consumption may not in any sense be caused by the change in wealth. Rather, both consumption and wealth simultaneously react to a shock, with shocks to productivity initially associated with either increases or decreases in consumption, depending upon whether or not such changes are anticipated; and, in the case of an anticipated future shock to productivity, consumption and wealth may move in opposite directions in the short term.

Theoretical and empirical approaches have succeeded in identifying various shocks that have differing implications for wealth and consumption, though the approaches appear to be still some way both from fully identifying the array of possible shocks affecting wealth and consumption, or in adequately explaining the path of wealth and consumption in recent years. There are other examples of shocks that may affect consumption and wealth to varying degrees. An increase in the average market power of firms for example, might lead to changes in equity prices and measured wealth, but consumption could fall if consumers are sensitive to any associated reductions in real disposable labour income arising from the effects of higher market power in reducing labour's expected share in total income. Similarly, structural changes that lead to greater economic stability and certainty (for example lower and less volatile inflation) may lead consumers to run down precautionary holdings of wealth and to increase consumption.

In short, the numerous causes of changes in wealth typically vary across countries and over time, and this might lead to correspondingly different responses of consumption and reduced-form estimates of the mpc_w. It appears plausible that country-specific structural shocks and structural change may have a marked impact on econometric estimates that is difficult to capture adequately in existing empirical models. The average mpc_w across time and countries based on reduced-form estimates might still be a useful summary statistic, albeit one around which there is wide uncertainty, particularly in samples limited to single countries and limited time periods.

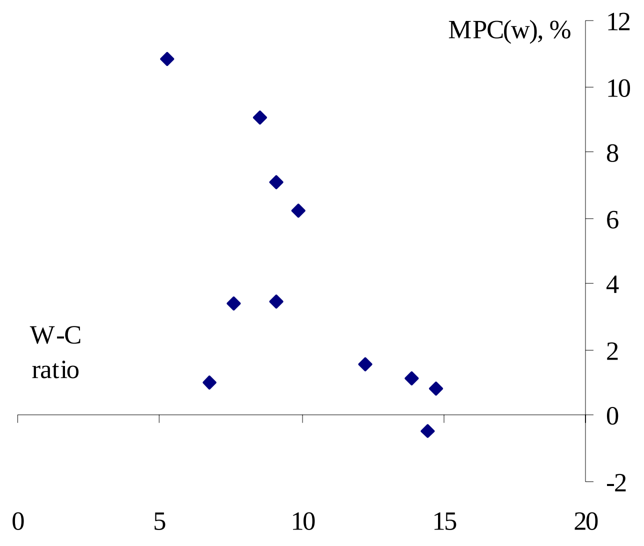
4.5 The role of data measurement, and other factors

As discussed in detail in Section 2, measurement differences may be important determinants of differences in estimates of mpc_w across countries. There have been wide differences in the measurement of financial wealth across countries, in particular differences in the form in which wealth is held (holdings of unquoted equity in France and some European countries is a case in point). There may be changes in the composition of wealth over time, some of which imply shifts from assets that were not previously included in the definition of financial wealth. For example, the definition of unquoted companies in Europe as discussed above includes plant and machinery of these companies as non-financial wealth. If such companies are floated then these become part of financial wealth. Persistent increases in the share of floated companies might therefore lead to a upward trending ratio of financial wealth to consumption.

The reduced-form, partial equilibrium approach to capturing the impact of changes in wealth on consumption faces a cocktail of such data problems and cannot account for underlying structural causes of simultaneous changes in both consumption and wealth. It is plausible, for example, that in circumstances when shocks to expected earnings dominate, a given shock to expected earnings is likely to have a similar impact on consumption in the two economies; but the response to measured changes in wealth could differ markedly. So in reduced-form estimates, the changes will not be captured consistently across countries. Economies where market capitalisation is low and wealth held in unquoted equities is *under*-recorded, might be (inaccurately) observed to have a *higher* mpc_w. This is because conventional empirical estimates of the mpc_w are commonly calculated by dividing an empirical estimate of the partial elasticity of consumption with respect to wealth by the observed wealth to consumption ratio. If the reaction of consumption to an expected earnings shocks is similar across countries, but the change in wealth is under-recorded because of data problems, then the mpc_w will be overestimated. We find evidence to support this view in Chart 9 where (using data taken from our SVAR estimates in Section 3.4) there is evidence of a negative relationship between the consumption-wealth ratio and the mpc_w. There is no economic reason why the apparent negative relationship should exist.

Moreover, even if data are measured correctly households' desired or target holding of financial wealth may vary across countries. One of the prime motivations for holding financial wealth is in order to smooth consumption in the face of variations in income. But government transfers,

Chart 9: MPCw and wealth-consumption ratios



including unemployment benefit and state-funded pensions, will influence this motivation: other things equal,[20] higher unemployment benefit and state-funded pensions will reduce the desired holding of financial wealth as they will smooth income and therefore consumption. So in countries where there is a relatively low state benefit provision, economic agents' target holdings of financial wealth may tend to be higher than in economies where there is relatively high state benefit provision: agents require a larger stock of personal wealth on which they can draw. Another reasons why desired holdings of wealth may differ is that households may be motivated to hold wealth as collateral, which they can use to gain access to capital markets on terms that are superior to those available without collateral. As such, differences in retail capital markets across countries may be one reason why households in different countries may desire to hold different amounts of wealth. Fleming (1973) considers capital market imperfections and whether they imply that consumption is constrained to current income, which will tend to increase the mpc_w.

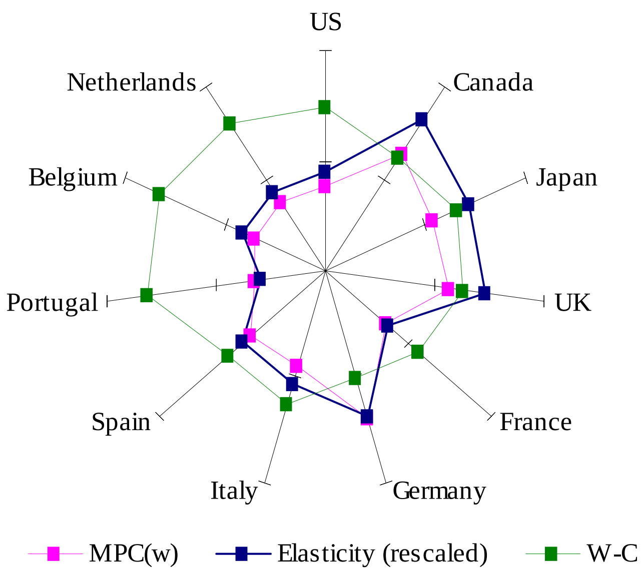
These differences in desired wealth holdings will be reflected in different wealth-consumption ratios across countries. Where the mpc_w is derived using a log-linear specification, they will in turn be reflected in differences in the mpc_w. Chart 10 shows the mpc_w, elasticity and wealth-consumption ratio for the eleven OECD countries; for each country, the elasticities (and therefore the mpc_w) shown are the average of the four SVAR estimates discussed in Section 3.4 and the wealth-consumption ratio is the average over the full sample period ie 1970 to 2002. The

(20) Pay-as-you-earn social security systems will also imply large future tax liabilities.

elasticities have been rescaled to enable comparison of the data points. It follows from equation **(9)** that, other things equal, a higher wealth-consumption ratio is associated with a lower mpc_w. This is borne out by Chart 10, which highlights the high wealth-consumption ratio/low mpc_w in the United States relative to Germany as discussed above.

Finally, as wealth-consumption ratios have varied over time (as shown in Charts 6 and 11), the use of a log-linear specification to calculate the mpc_w means that it is making a moving target in empirical estimation within and across countries.

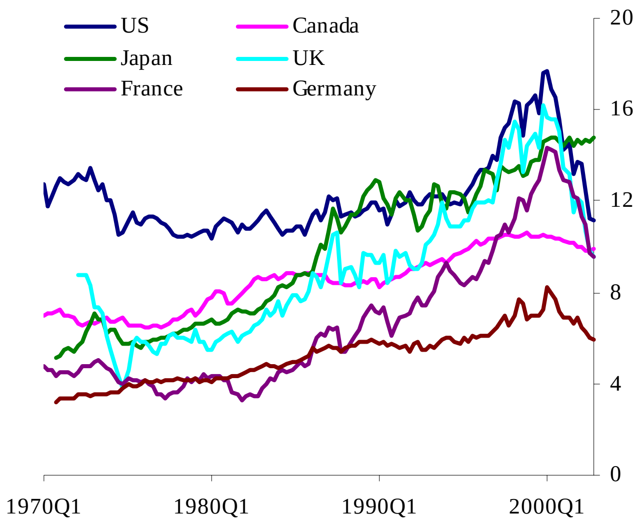
Chart 10: MPCw, elasticities and wealth-consumption ratios by country



4.6 Implications for cross-country comparisons

Our analysis has suggested that there is little theoretical rationale for the wide cross-country dispersion of empirical estimates reported in the literature. We find it unlikely that differences in economic factors such as the length of planning horizon, the rate of return on financial assets and in the form of wealth holding are large enough to explain the divergences in cross-country differences in mpc_w reported in the literature. Some of the divergence could be explained by housing, even though we found this consideration to yield no robust results in our econometric work (this may be explained by measurement difficulties which are even more acute than for other wealth measures): to the extent there is any short-run link between the housing market and consumption, part of the explanation may be related to strong growth in some countries' housing

Chart 11: Wealth-consumption ratios over time



markets at certain times in the sample period. But these short-run effects should have little long-run impact on consumption so appear unlikely to have a marked impact on long-run estimates of the mpc_w. As such, the differences in empirical estimates may reflect factors other than demographic and structural differences across countries, such as inconsistencies in the measurement of wealth across countries and a failure to account for the underlying structural causes of changes in both consumption and wealth.

Overall, and in spite of the apparently large differences in point estimates of the mpc_w across countries, we may not be able to reject the hypothesis that such differences are significant. If measurement difficulties are particularly important, for example, apparently large differences may be associated with imprecise point estimates. This is an issue we explore further in the following section.

5 A dynamic panel approach to the estimation of mpc_w

Since the long-run mpc_w should be similar across countries if planning horizons and rates of return are similar, it is sensible to exploit this at the estimation stage. One way to do so is to use a panel specification in which we can test the null hypothesis of cross-sectional homogeneity. However, since the short-run mpc_w is more likely to vary across countries, the panel framework needs to be dynamic and allow for heterogeneity in the short-term deviations from the long-run

mpc_w. Recent contributions have attempted to tackle this by using traditional panel techniques. Bayoumi and Edison (2003), for example, adopt a two-step procedure, estimating first a long-run equilibrium relationship between consumption and wealth, and then using the estimated expression as an error correction term in a standard dynamic consumption function. However, their approach implies that the hypothesis that the mpc_w is identical across countries cannot be tested formally. This appears to be a limitation, especially as they investigate whether there is a difference in the mpc_w between countries with bank-based and market-based financial systems.

In this paper, we seek to address these concerns by estimating the mpc_w using a ratio specification, hence avoiding the use of volatile wealth-consumption ratios, and by adopting a dynamic panel framework which can handle both long-run homogeneity and short-run heterogeneity of the parameters. A suitable framework has been suggested recently by Pesaran, Shin and Smith (1999).[(21)] Subject to a specification in which the parameters can be interpreted as mpc_w, their pooled mean group estimator (PMGE) provides a means to impose a (testable) cross-country restriction on the long-run mpc_w, while taking into account possible cross-country differences in the adjustment to this common long-run mpc_w. A PMGE-based estimate of the mpc_w therefore is likely to provide a theory-consistent guide to the wealth effects on consumption, and a better guide than estimates obtained from either traditional panel methods or single equation techniques.

Moreover, due to its flexibility, the PMGE is particularly suitable for the 'large N, large T' case, ie when cross-sectional and time-series dimension are both large and similar in magnitude, as is the usual case in cross-country studies.[(22)] In this case, traditional panel methods, including the fixed and random effects estimators, are overly restrictive, since they force all coefficients except the intercepts to be identical across countries. By contrast the mean group estimator (MGE), which is the average coefficient of separate regressions for the cross-sectional units, is not restrictive enough since it does not impose any cross-sectional restrictions at all. The PMGE is an intermediate procedure combining both pooling and averaging. Moreover, as shown by Pesaran *et al* (1999) the PMGE estimator is asymptotically normal for the cases of both stationary and non-stationary regressors, so that regardless of the time-series properties of the data the parameters can be estimated by maximum likelihood methods using standard optimisation algorithms.

(21) We are grateful to Hashem Pesaran for having made available code for this framework.
(22) Note that there are 11 countries in our panel, so that N=11, and we use quarterly data from 1970 Q1 to 2002 Q4, so that T=134, so even though in our application $N<T$, our panel is clearly of the type envisaged by Pesaran *et al* (1999) as neither N nor T are very large.

For the panel equations, we choose a specification in which the cointegration relationship equals **(6)**, and so provides an estimate of the long-run mpc_w which needs not to be backed out from an estimate of the elasticity, so that the restriction on the mpc_w can be imposed directly. This specification is

$$\Delta \left(\frac{C_t}{Y_t}\right)_i = \gamma_{0,i} + \gamma_{1,i} \Delta \left(\frac{C_t}{Y_t}\right)_i + \gamma_{2i} \Delta \left(\frac{W_t}{Y_t}\right)_i + \alpha_i \left[\frac{C_t}{Y_t} + \beta \frac{W_t}{Y_t}\right]_i + \varepsilon_t \qquad \textbf{(16)}$$

where $\frac{C_t}{Y_t}$ is a consumption-income ratio, $\frac{W_t}{Y_t}$ is a wealth-income ratio and the coefficient β provides the estimate of the long-run mpc_w.[23] The coefficient β is constrained across countries, while the adjustment coefficient α, the short-term coefficients γ, as well as the error variances are estimated separately for each country, as indicated by the index i. As discussed in Palumbo *et al* (2002), if there is a stable long-run relationship between the consumption-income and wealth-income ratios, the error correction term in **(16)** would be stationary, and for the series used in this paper the panel evidence we present in the annex seems to support this.

5.1 Benchmark estimates of the mpc_w

In order to obtain a benchmark estimate, we use the broadest measures of C_t and W_t: total private sector consumption and net financial wealth, and assume that there is a long-run equilibrium relationship linking these variables when expressed as a ratio to disposable income. We have also estimated the panel using the two alternative income measures (disposable income adjusted for net taxes, as well as labour income, which excludes all income from assets) and with smaller panels, eg for the G7 only. While we find that for disposable income and labour income the results are robust to the choice of the whole panel or the G7, this is not always the case when we use the alternative income measures.

Table D shows the results for our benchmark, in terms of the point estimates and corresponding standard errors, as well as standard diagnostic tests on serial correlation, functional form, normality and heteroscedasticity. The point estimates are quoted on the basis of quarterly data, so need to be multiplied by 4 to obtain standard mpc_w as quoted in the literature. Our benchmark estimate therefore, based on the PMGE estimator, is 1.7, or 6.8 at an annualised basis.[24] This estimate is highly significant, with a t-ratio of about 10. Table D also reports the MGE estimator,

(23) Specifications of this type have also been estimated by Davis and Palumbo (2001).

(24) Using the theoretically preferred income measure (disposable income adjusted for net taxes, as well as labour income, which excludes all income from assets) we find an estimate of 0.6, or 2.4 at an annualised basis. Using the labour income measure, we find an estimate of 0.45, or 1.8 at an annualised basis.

Table D: Benchmark estimates of long-run mpc_w

Dynamic panels	Estimate (β)	Standard error (σ_β)	LR	p-value	H	p-value
PMGE	0.017	0.001	44.73	0.00	0.62	0.43
MGE	0.012	0.006	..	..	..	..
Dynamic OLS	Estimate (β)	Standard error (σ_β)	SC	FF	NO	HE
Belgium	-0.004	0.006	1.13	0.88	1.62	1.79
Canada	0.066	0.048	0.20	0.10	1.05	0.27
France	-0.002	0.004	0.02	0.86	4.10	0.39
Germany	0.008	0.002	0.01	0.88	2.91	0.11
Italy	0.024	0.006	3.05	2.06	0.02	0.23
Japan	0.018	0.002	0.08	3.23	72.2	1.88
Netherlands	0.001	0.004	9.83	0.58	13.01	1.51
Portugal	-0.004	0.001	0.71	0.16	0.63	0.04
Spain	-0.002	0.001	3.17	0.34	121.57	2.04
UK	0.017	0.013	1.97	0.03	1.51	1.14
US	0.014	0.003	0.21	0.75	4.89	1.82

Abbreviations: PMGE pooled mean group estimator, MGE mean group estimator, LR likelihood ratio test for long-run homogeneity, H Hausman test for long-run homogeneity, SC test for serial correlation, FF test for functional form, NO test for normality, HE test for heteroscedasticity.

as well as the OLS regressions for the individual countries from which it is derived. This estimate of the long-run mpc_w is somewhat lower at 1.2, or 4.8 at an annualised basis, and significant at conventional levels.[25]

When we use the appropriate test statistic to test for our hypothesis of a common equilibrium relationship and common long-run mpc_w for the countries in our panel, we find that the null hypothesis of a common long-run cannot be rejected: the Hausman test has a statistic of 0.62, with a p-value of 0.43.[26]

(25) Even the estimates presented in this section may be subject to the caveats identified in the paper, eg difficulties in the measurement of wealth across countries and a failure to account for the shocks causing changes in both consumption and wealth.

(26) The Hausman test is based on the difference between the MG estimator, which is consistent but not efficient under the null of homogeneity, and the PMG estimator, which is both consistent and efficient. The test statistic is

$$H = \widehat{\beta}'[var(\widehat{\beta})]^{-1}\widehat{\beta} \quad \sim\chi^2(n),$$

where $\widehat{\beta} = \beta_{PMG} - \beta_{MG}$ is a vector of differences between the MG and PMG estimators (the statistic has a χ^2 distribution with n degrees of freedom, where n is the number of cross-sectional units). Under the null hypothesis, the term $\widehat{\beta}$ will be small and the term $[var(\widehat{\beta})]^{-1}$ will be large, so that it will not be rejected. If T is sufficiently large, which is not the case in our application, the null hypothesis of homogeneity could also be tested using the relative likelihood of the MG and PMG models. The LR test statistic is

$$LR = 2(L_{PMG} - L_{MG}) \quad \sim\chi^2((n-1)*k),$$

where k is the number of stochastic parameters. Under the null hypthesis, the term $L_{PMG} - L_{MG}$ will be small and so it will not be rejected. However, if T is small, the likelihood ratio test will tend to reject the null even if it is true, and therefore should not be used (this is the case in our application).

For the individual countries, the estimates are quite mixed. Not imposing homogeneity therefore would lead to quite different results across countries, and some of these results clearly do not make much economic sense - notably the negative mpc_w estimated for Belgium, France, Portugal and Spain, as they would suggest that a rise in wealth leads to a fall in consumption in the long run. Pesaran *et al* (1999) have argued that this is one of the situations in which the PMG estimator may be particularly useful.[27]

5.2 Short-run adjustment to the long-run mpc_w

Having discussed the estimates of the long-run mpc_w, for which the PMGE estimator imposes cross-country homogeneity, we now take a closer look at the estimated adjustment coefficients, which are not constrained by the PMGE estimator. Table E shows these coefficients using both the PMGE and MGE. The first observation from looking at Table E is that there are considerable differences between the two sets of estimates, ie imposing long-run homogeneity leads to different estimates of the short-run dynamics. Using the PMGE, the estimates for the adjustment coefficient are generally smaller. Moreover, they fall into a smaller range than using the MGE, suggesting that differences in the speed of adjustment across countries may be overstated when using single-country regressions.

Table E: Estimates of the adjustment coefficients

	PMGE		MGE	
	Estimate (α)	Standard error (σ_α)	Estimate (α)	Standard error (σ_α)
Belgium	-0.0008	0.0093	-0.0343	0.0200
Canada	-0.0094	0.0275	-0.0287	0.0291
France	0.0020	0.0142	-0.0670	0.0367
Germany	-0.1028	0.0409	-0.2183	0.0639
Italy	-0.0195	0.0144	-0.0330	0.0222
Japan	-0.1734	0.0527	-0.1706	0.0535
Netherlands	-0.0127	0.0186	-0.0811	0.0549
Portugal	-0.0119	0.0104	-0.2268	0.0720
Spain	-0.0041	0.0077	-0.1682	0.0474
UK	-0.0756	0.0334	-0.0768	0.0534
US	-0.1114	0.0431	-0.1380	0.0504

Abbreviations: PMGE pooled mean group estimator, MGE mean group estimator.

The speed of adjustment is highest in Germany, Japan, the United Kingdom and the United States,

(27) In fact, eliminating these countries from the panel yields basically the same estimate for the mpc_w as for the full panel. The Hausman test in this case has a statistic of 0.24 and a p-value of 0.63.

where the coefficients are larger than 0.1 in absolute value. At the other end of the spectrum, the adjustment is slowest in Belgium, Canada and Spain, which have a coefficient of less than 0.01 in absolute value, with the remaining countries in the panel in between.[(28)] In general, the adjustment coefficients are quite low. So while we find that imposing a common long-run parameter is not rejected by the Hausman test and yields a sensible estimate, one caveat is that the low degree of error correction suggests that there is weak evidence for cointegration in equation **(16)**.

5.3 Robustness checks

Although the benchmark estimate is reasonable, it is necessary to check for robustness. We do this in two ways. First, we assume that wealth effects on consumption are non-linear, ie there are differences between the effects of small and large changes in wealth. Second, we allow for asymmetries depending on the sign of the change in wealth, ie a difference between the effects of increases and decreases in wealth. In both cases, we compare results to the benchmark estimate presented above.

Table F: Non-linearities and the mpc_w

Dynamic panels	Est. (β)	Std.err.(σ_β)	Est. (β^*)	Std.err. (σ_{β^*})	LR	p-value	H	p-value
PMGE	0.036	0.004	-0.001	0.000	65.36	0.00	..	..
MGE	0.034	0.018	-0.001	0.001	..	..	..	..
Dynamic OLS	**Est. (β)**	**Std.err.(σ_β)**	**Est. (β^*)**	**Std.err. (σ_{β^*})**	**SC**	**FF**	**NO**	**HE**
Belgium	0.005	0.042	0.000	0.002	0.28	0,86	2.42	1.48
Canada	0.209	0.677	-0.009	0.040	0.31	0.14	0.85	0.27
France	-0.004	0.021	0.000	0.001	0.03	0.52	2.82	0.98
Germany	0.012	0.020	0.000	0.002	0.28	4.08	3.18	0.99
Italy	0.024	0.026	0.000	0.002	17.20	3.54	0.19	1.62
Japan	0.029	0.011	-0.001	0.001	0.08	1.78	68.93	2.12
Netherlands	0.012	0.012	0.000	0.000	0.05	0.74	46.43	0.00
Portugal	0.009	0.008	0.000	0.000	0.50	0.11	0.57	0.00
Spain	0.004	0.002	0.000	0.000	1.89	0.23	227.62	0.59
UK	0.029	0.109	-0.001	0.005	1.01	0.00	2.33	0.08
US	0.043	0.038	-0.001	0.002	1.64	0.30	9.85	1.42

Abbreviations: PMGE pooled mean group estimator, MGE mean group estimator, LR likelihood ratio test for long-run homogeneity, H Hausman test for long-run homogeneity, SC test for serial correlation, FF test for functional form, NO test for normality, HE test for heteroscedasticity.

(28) Note that we actually obtain a positive estimate of the adjustment coefficient for France.

5.3.1 Non-linearities and the mpc_w

A simple way to allow for non-linearities is to include a quadratic term in equation **(16)**. If non-linearities are important, this term should be significant, and the estimate of the mpc_w should be similar to the benchmark estimate. We prefer a quadratic term to using a dummy-interacted term based on and (necessarily) arbitrary threshold. Our results are shown in Table F. They suggest that non-linearities are statistically significant. The point estimate of the coefficient on the term capturing the non-linearities is negative, suggesting that large changes in wealth have proportionately smaller effects on consumption, perhaps because they are more likely to be reversed than smaller changes. The inclusion of this term also leads to a larger point estimate for the long-run mpc_w, which increases to 3.6 or 14.4 at an annualised basis, which is twice the benchmark estimate. Looking at the individual country regressions, we see that the point estimates on the additional coefficient are of the same magnitude and significance for most countries. For Canada, the point estimate is remarkably high, but not statistically significant. We also see that the estimate for the long-run mpc_w generally increases, and that the negative estimates from our benchmark specification turn positive, except in the case of France.

Table G: Asymmetries in the mpc_w

Dynamic panels	Est. (β)	Std.err.(σ_β)	Est.(β^*)	Std.err.(σ_{β^*})	LR	p-value	H	p-value
PMGE	0.018	0.002	-0.004	0.001	..	..	..	..
MGE	0.011	0.004	0.007	0.008	..	..	..	..
Dynamic OLS	**Est. (β)**	**Std.err.(σ_β)**	**Est.(β^*)**	**Std.err.(σ_{β^*})**	**SC**	**FF**	**NO**	**HE**
Belgium	-0.003	0.005	0.000	0.002	1.04	1.04	2.28	1.84
Canada	0.036	0.039	0.086	0.129	0.23	0.11	0.90	0.42
France	-0.006	0.007	0.006	0.007	0.73	0.50	4.06	2.98
Germany	0.010	0.003	-0.003	0.002	0.19	1.20	3.18	0.06
Italy	0.024	0.004	0.002	0.004	22.39	0.91	1.36	0.31
Japan	0.019	0.002	-0.002	0.001	0.58	4.52	59.45	2.13
Netherlands	-0.001	0.003	0.001	0.002	0.14	0.73	33.90	0.01
Portugal	-0.002	0.001	-0.004	0.003	0.79	0.00	1.54	0.90
Spain	-0.002	0.001	-0.002	0.002	2.97	0.02	106.53	1.37
UK	0.028	0.030	-0.009	0.002	2.13	0.29	0.55	1.79
US	0.014	0.003	-0.002	0.001	0.22	0.13	6.25	1.80

Abbreviations: PMGE pooled mean group estimator, MGE mean group estimator, LR likelihood ratio test for long-run homogeneity, H Hausman test for long-run homogeneity, SC test for serial correlation, FF test for functional form, NO test for normality, HE test for heteroscedasticity.

5.3.2 *Asymmetries and the mpc_w*

In order to capture potential asymmetries, we include a variable interacting the wealth to income ratio with a binary dummy taking a value of zero if the wealth to income ratio has increased in period t.[29] This means that the sum of the two coefficients now provides an indication of the long-run mpc_w for periods in which the wealth to income ratio has fallen, while the estimate of the existing coefficient now relates only to periods in which the wealth to income ratio has increased. The results are shown in Table G. The estimate for β is 1.8, or 7.2 in annualised terms, which suggests that in periods of rising wealth-consumption ratios, the mpc_w is larger than in the benchmark case. The interacted variable has a negative coefficient of -0.4, or -1.6 in annualised terms, which means that the mpc_w is 1.4, or 5.6 on an annualised basis for periods in which the wealth to income ratio has fallen. The fact that the former is closer to the benchmark estimate seems to reflect the fact that wealth to income ratios have generally increased over the sample.

6 Conclusions

This paper offers a critique of the existing literature on the wealth effect on consumption. At the outset, we have showed that forecasters have been relatively unsuccessful in predicting the cross-country strength of the effect of changes in equities, and financial wealth more generally, on consumption. Our analysis suggests there are several reasons why this is the case, and in terms of its conclusions this paper departs from the previous literature.

Our assessment is that wide cross-country differences in empirical estimates of the mpc_w in individual countries, including new evidence produced in this study, when taken at face value, are potentially misleading. Such differences are not useful for policymaking unless they are rooted in stable and explainable structural differences. Yet our analysis suggests that cross-country differences in rates of return and planning horizons can explain only a small part of the variance in empirical cross-country estimates of the mpc_w. Furthermore, our analysis of demographic factors gives little support to the view that the mpc_w is relatively high in the United States, even though this is a feature of both empirical and theoretical studies considered in this paper. There is also little evidence that cross-country differences in the mpc_w can be attributed to the structure of asset portfolios, including the relative importance of indirect versus directly held equities. Conversely,

(29) Given our timing convention, this implies a decrease in the wealth to income ratio at the beginning of period t relative to the beginning of period $t-1$.

there are reasons to believe that the mpc_w may vary systematically over short and medium-term horizons following shocks that change house prices, but data deficiencies prevent us from testing this hypothesis directly.

Measures of wealth that closely proxy expectations of future income are difficult to conceptualise, still more difficult to measure. We argue that much of the differences in empirically estimated mpc_w are attributable to data deficiencies. In our view, partial equilibrium approaches to capturing the impact of changes in wealth on consumption face a cocktail of data problems and a failure to account for underlying structural causes of simultaneous changes in both consumption and wealth. This part of our conclusions is slightly negative, perhaps even discouraging, to international forecasters.

A more positive implication of our analysis is that, when there are grounds for believing that an international shock is approximately symmetric across countries (for example to expected earnings), the impact on consumption might be expected to be more similar across countries than most international forecasting models would predict. As such, it is plausible that on average, across assets, countries and time, the mpc_w remains a useful summary statistic illustrating the average response of wealth and consumption to various structural shocks. When we use a suitable panel technique we find that the hypothesis of the long-run mpc_w being the same across countries cannot be consistently rejected, and obtain a plausible estimate for the cross-section of eleven OECD countries. This estimate is a little over 6%, broadly consistent with estimates used in a wide range of policy models.

Annex

Table H: Unit root tests for consumption

	Total consumption (log)				Durables consumption (log)			
ADF tests (incl trend)	Lags	Level	Lags	Diffs	Lags	Level	Lags	Diffs
Belgium	0	*-3.34	1	***-10.19	0	*-3.17	0	***-518.23
Canada	0	*-3.25	0	***-110.49	0	**-3.48	0	***-126.91
France	0	*-3.41	3	***-4.44	3	-3.04	2	***-6.81
Germany	1	-2.29	0	***-74.84	1	-2.23	0	***-77.91
Italy	1	-1.76	0	***-135.12	1	-1.71	0	***-144.63
Japan	2	-0.83	1	***-16.11	2	-1.43	1	***-16.13
Netherlands	5	-3.07	4	**-3.78	5	*-3.26	4	**-3.70
Portugal	2	**-3.47	1	***-6.80	2	*-3.25	1	***-6.85
Spain	2	*-3.28	1	***-5.16	2	-2.99	1	***-5.71
UK	0	-1.21	0	***-79.56	0	-0.94	0	***-96.41
US	3	**-3.95	0	***-153.29	3	**-4.57	3	**-3.68
Panel tests								
Im Pesaran Shin	..	***-2.33	..	***-223.16	..	***-2.41	..	***-380.76

The table shows test statistics and significance levels for the null hypothesis of a unit root (* 10%, ** 5%, *** 1%). The number of lags in the ADF regressions is determined using the Schwarz Bayesian information criterion. The panel test is based on an average of the country ADF tests (for details, see Im, Pesaran and Shin (2003)).

Table I: Unit root tests for income

	Disposable income (log)				Labour income (log)			
ADF tests (incl trend)	Lags	Level	Lags	Diffs	Lags	Level	Lags	Diffs
Belgium	8	-2.78	2	***-4.97	0	*-3.24	4	-2.98
Canada	0	-2.99	0	***-13.27	2	*-3.30	2	***-4.83
France	5	-3.04	4	**-3.93	6	*-3.28	8	-2.14
Germany	0	-2.52	0	***-75.30	0	*-3.28	0	***-173.92
Italy	7	-1.62	6	**-3.51	1	-2.99	0	***-268.22
Japan	2	-2.12	1	***-10.38	1	-2.63	4	*-3.29
Netherlands	6	-1.88	5	**-3.82	5	-2.76	4	-2.16
Portugal	8	*-3.35	7	**-3.73	8	-2.25	6	**-3.88
Spain	5	-2.38	4	**-3.70	8	**-3.46	8	-2.20
UK	1	-2.05	0	***-13.01	2	-1.65	2	**-3.77
US	0	*-3.19	0	***-105.72	2	*-3.37	1	***-10.08
Panel tests								
Im Pesaran Shin	..	**-1.67	..	***-83.39	..	***-3.28	..	***-172.78

The table shows test statistics and significance levels for the null hypothesis of a unit root (* 10%, ** 5%, *** 1%). The number of lags in the ADF regressions is determined using the Schwarz Bayesian information criterion. The panel test is based on an average of the country ADF tests (for details, see Im *et al* (2003)).

Table J: Unit root tests for wealth

	Total financial wealth (log)			
ADF tests (incl trend)	Lags	Level	Lags	Diffs
Belgium	1	-1.89	4	-2.89
Canada	9	-1.41	0	***-215.41
France	1	-2.25	8	-2.70
Germany	0	-1.11	4	-1.97
Italy	4	-2.45	3	**-4.02
Japan	3	-1.71	4	-2.82
Netherlands	0	-2.54	4	-1.31
Portugal	1	-2.47	8	-2.99
Spain	2	-1.66	4	-2.82
UK	0	-3.08	4	-2.74
US	0	-1.74	3	-2.72
Panel tests				
Im Pesaran Shin	..	-0.25	..	***-82.97

The table shows test statistics and significance levels for the null hypothesis of a unit root (* 10%, ** 5%, *** 1%). The number of lags in the ADF regressions is determined using the Schwarz Bayesian information criterion. The panel test is based on an average of the country ADF tests (for details, see Im *et al* (2003)).

Table K: Unit root tests for wealth

	Equity wealth (log)				Non-equity wealth (log)			
ADF tests (incl trend)	Lags	Level	Lags	Diffs	Lags	Level	Lags	Diffs
Belgium	0	-1.85	0	***-4.97	0	***-7.30	1	***-12.56
Canada	5	-1.53	8	***-13.28	8	**-3.48	7	-2.62
France	1	-2.85	0	**-3.93	0	**-3.56	1	***-9.67
Germany	0	-1.00	0	***-75.30	0	*-3.36	1	***-10.08
Italy	4	-2.73	6	**-3.51	3	-0.81	2	***-13.39
Japan	0	0.46	1	***-10.38	2	-2.25	1	***-11.86
Netherlands	3	*-3.34	5	**-3.82	0	***-7.04	0	***-16.11
Portugal	1	-1.91	7	**-3.73	1	-2.16	0	***-13.66
Spain	3	-2.13	4	**-3.70	2	-1.46	2	***-8.79
UK	0	-1.76	0	***-13.01	0	***-7.56	2	***-8.21
US	0	-2.04	0	***-105.72	7	2.78	1	***-15.38
Panel tests								
Im Pesaran Shin	..	1.21	..	***-83.39	..	***-4.78	..	***-38.00

The table shows test statistics and significance levels for the null hypothesis of a unit root (* 10%, ** 5%, *** 1%). The number of lags in the ADF regressions is determined using the Schwarz Bayesian information criterion. The panel test is based on an average of the country ADF tests (for details, see Im *et al* (2003)).

Table L: Cointegration tests

	Test statistic
Pedroni panel statistics	
v	***5.51
rho	***-3.26
pp	***-2.62
adf	-0.63
Pedroni group statistics	
rho	***-4.01
pp	***-3.29
adf	-0.95

The table shows the test statistics and significance levels for the null hypothesis of a unit root in the residuals, ie no cointegration, of consumption, income and wealth (* 10%, ** 5%, *** 1%). The panel statistics assume that the properties of the residuals are identical across countries, the group statistics make no such assumption. All statistics are based on Pedroni (1999).

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