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Consumption, personal income, financial wealth, housing wealth, and long-term interest rates:

A panel cointegration approach for 50 US states

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Abstract

This study investigates the long-run and short-run relationship between consumption, income, financial and housing wealth, and a long-term interest rate for the 50 US states. Using an updated set of quarterly data from 1975 to 2018, we perform panel cointegration analysis allowing for cross-sectional dependence. We obtain the following results. First, there is strong evidence for cointegration among consumption and its determinants. Second, estimates of the housing wealth and financial wealth elasticity of consumption range from 0.072 to 0.115 and 0.044 to 0.080, respectively. Finally, Granger causality tests show that there is a bidirectional short-term causality between per capita consumption, income, and financial wealth in the short run and between all the variables in the long run.

JEL codes: E21; E44; R31.

Keywords: Consumption; Financial Wealth; Housing Wealth; Wealth effects; 10-year Treasury Constant Maturity Rate; Panel Cointegration; Granger Causality.

1 Introduction

During recent years, the consumption response to personal income, housing and financial wealth changes has received substantial attention by policymakers and economists. Increases in house prices associated with financial innovations from 2002 to 2009 led the US economy to the most severe crisis since the post-World War II era. This crisis spread a deep global recession from which some countries have only recently recovered. Some observers suggest that monetary policy played a central role in the crisis, while others support the position that policy was appropriate under the macroeconomic circumstances at that time and the information that was available to policymakers. In response to the US crisis, on December 16, 2008, the U.S. Federal Reserve's Federal Open Market Committee (FOMC) lowered its traditional monetary policy instrument (federal funds rate) to zero but the recovery remained weak. New unconventional monetary policy tools, such as quantitative easing (QE) and forward guidance, were introduced by the FOMC in response to the inability of the traditional low interest rate policy to stimulate the economy arising from the zero-bound problem.

At the same time, given the deficiency of the federal funds rate as an effective monetary policy instrument, economists proposed new methods to identify unconventional monetary policy and estimate its effects on the real economy. According to the most recent literature (Greenlaw, Hamilton, Harris, and West 2018; Swanson 2015; Williams 2013; Wright 2012; Balduzzi, Elton, and Green 1996), the 10-year treasury constant-maturity rate is strongly associated with QE and is a factor in the market's response to the announcements of the FOMC. Therefore, we conjecture that this interest rate plays an important role in controlling consumption. In the macroeconomics literature, the subject of the determinants of consumption remains a vital open issue for discussion among policymakers, researchers and observers. Therefore, the objective of this paper is to estimate a Keynesian consumption equation that incorporates both wealth and interest rate effects.

In this study we investigate the consumption behaviour in the 51 US states. The theoretical literature reveals that the primary determinants of consumption are income, financial wealth, housing wealth and interest rates. Keynes (1936) suggests that consumption mainly depends on income for a given level of employment and additional factors that should influence spending. Recent literature uses different methodologies and data and concentrates on wealth types that should influence spending (Case

Quigley, and Shiller 2005, 2011, 2013; Carroll, Otsuka, and Slacalek 2011; Slacalek 2009 etc.).

The main motivation of this work is understanding wealth effects from another view. This is done by adding the 10-year treasury constant maturity rate in the consumption equation and using a new data set based mostly on housing behaviour. We can then shed light on the effect of the interest rate on consumption which from a theoretical point of view is ambiguous, as it depends on the interplay of substitution and income effects (Romer, 2012). The second piece of motivation arises from the importance of the long-term interest rate as a determinant of spending in unconventional monetary policy. This factor has been suggested by the literature as a tool for monetary policy (Balduzzi et al. 1996; Swanson 2015). In this way, we can examine whether the model can explain the effects of unconventional monetary policy on consumption. This is done in two ways. First, we use the methodology of panel cointegration and error-correction proposed by Pedroni (2004) and Westerlund (2007) to test for the long-run relationship between consumption and its four determinants: income, financial wealth, housing wealth, and the interest rate. Second, we employ two Granger causality techniques to address the short-run causal relationships among the variables.

This study contributes to the related literature along the following lines. First, we use a newly calculated and updated data set from 1975 to 2018. Second, we test for a long-run relationship among consumption and its determinants using a panel data analysis for the 50 US states over 43 years. Third, we include in the analysis the long-term interest rate as an important variable that relates monetary policy with the real economy, following the unconventional types of monetary policy applied in the US since the onset of the financial crisis. Fourth, we use econometric techniques that allow us to address the endogeneity across variables, heterogeneity, and cross-sectional dependence in the panel. Fifth, we investigate the short run and long-run causal relationships among the variables of interest.

The main results of the study are as follows: First, we find strong evidence for a long-run equilibrium relationship among consumption, income, financial and housing wealth, and a long-term interest rate. Second, in the long run, we find strong evidence for positive income and housing wealth effects on consumption. The estimated income elasticity varies between 0.62 and 1.08, depending on the model specification and the estimated housing wealth elasticity varies between 0.07 and 0.12. Financial wealth and the long-term interest rate do not seem to affect consumption significantly in the long run. Second, in contrast to the long-run effects, in the short run, both forms of wealth and the interest rate are also significant and have a positive and negative effect on consumption, respectively. However, the effect of housing wealth seems to be larger than that of financial wealth. Finally, short-run and long-run causality tests show evidence for the linkage of housing wealth with the 10-year treasury rate in most specifications, confirming that developments in the housing market have a major and predictable effect on real economic activity.

The rest of this study is organized as follows. Section 2 reviews the literature on the determinants of consumption. Section 3 describes the data and our model. Section 4 presents our empirical methodology and findings. Section 5 discusses the main results and Section 6 concludes.

2 Literature review

There is an extensive literature investigating the wealth effects on consumption for different countries and sample periods and employing various methodologies. The literature concludes that the long run marginal propensity to consume (MPC) out of housing wealth is about 5-10 cents per dollar. However, MPC depends crucially on the institutional characteristics of each country and the type of data, suggesting that MPC is much larger in developed financial markets and in the presence of liquidity constraints (Iacoviello 2011; Sierminska and Takhtamanova 2012; Case et al. 2013; Carroll, Slacalek, Tokuoka, and White 2017). In contrast to countries with no market-based economy, the effect of housing wealth is larger than that of financial wealth for the U.S and UK (Slacalek 2009). Equally, the housing wealth effect has risen substantially after 1988 as it has become easier to borrow against housing wealth, especially in countries with more developed mortgage markets (Case et al. 2013; Slacalek 2009; etc.).

Sierminska et al. (2012) in a sample that includes Canada, Finland, Italy, Germany, and US estimate consumption elasticity with respect to financial wealth to range from 0.01 to 0.03, while the elasticity for housing wealth is 0.04. Campbell and Cocco (2007) suggest that in the UK, elasticity of consumption with respect to housing wealth is 1.22. Correspondingly, Atalay, Whelan, and Yates (2016) using aggregate data suggest that MPC out of housing wealth is around 0.02-0.03 for Australia and Canada by. Hori and Niizeki (2019) investigate the housing wealth effects in Japan over the period 1983-2012. The results report that the consumption response to an increase in housing wealth ranges from 0.0059 to 0.0082 and appears to be larger for older households than that of younger households. While in China the estimated elasticity of consumption for housing wealth is as high as 0.19, much larger in magnitude (almost 10 times) than that of financial wealth (Chen, Hardin, and Hu, 2018).

Case et al. (2012) investigate the effects of wealth on consumption by using quarterly state level data from 1975 to 2012 and apply multiple specifications. They introduce the error correction model and find a co-integrated relation between consumption and income. Under the assumption that income includes that derived from the stock market and housing, they don't apply the cointegration method on housing and financial wealth. It is assumed that log levels of consumption and income are cointegrated with the known cointegrating vector (1, -1), suggested by previous research Park (1992) and King, Plosser, Stock, and Watson (1991).

Han and Ogaki (1997) examine the post war consumption and income in US by employing two different methodologies to test cointegration. The canonical cointegration regression suggested by Park (1992) imposing restriction only in the stochastic trend and the cointegration methodology based on stationarity of the difference of log consumption and log income. The results indicate that there is no cointegration between the variables.

More recent studies extend the model of PIH by considering the different types of wealth. Carroll et al. (2011) estimate the eventual MPC out of financial and housing wealth to be in the ranges 0.041-0.064 and 0.087-0.159, respectively. Finally, in US, housing wealth is 7.04 cents against financial which is about 5 cents. After 1988, although financial wealth remains unchanged around 3-4 cents, the housing wealth effect grew from almost zero to 3 cents due to mortgage markets development and their link to housing wealth in most countries. The results are in line with Iacoviello (2011) that the co-movement between housing wealth and consumption is larger in developed financial markets and in the presence of liquidity constraints. Especially in US, the passage of Tax Reform Act loosened borrowing constraints and facilitated borrowers to use the housing equity for consumption (Case et al. 2012; Slacalek 2009). Higher house prices can have substantial effects on consumption as they raise borrowing capacity (Case et al. 2012; Slacalek 2009; Aladangady 2017). Rapach and Strauss (2006) suggest that the housing wealth effect is not uniform across the Eighth District states. Based on cointegration methods, they support the existence of a stable long-run relationship between personal consumption, personal income, housing and financial wealth.

Angrisani, Hurd, and Rohwedder (2019) focus on a sample of American adults over the age of 50 because they are more likely to be home owners and respond distinctly to home values. They provide micro evidence for MPC out of an unexpected change in housing wealth to be 6 cents per dollar for middle-aged consumers.

Generally, households with different levels of wealth have different estimated elasticities of consumption (Calomiris, Longhofer, and Miles 2013; Bampinas et al. 2017). Bampinas, Konstantinou, and Panagiotidis (2017) provide evidence for the long-run coefficient of wealth in the range of 0.053-0.088 by considering the role of inequality and demographics for the 48 US states. Christou, Gupta, Nyakabawo, and Wohar (2018) investigate the long run relation of housing wealth and non-housing consumption during the period 1953-2016 by using various quantile cointegration methods. They conclude that the variables are at most cointegrated at lower quantiles. Additionally, low house prices act as an inflation hedge when inflation is relatively large.

In the same line, Vinson (2018) used panel data for US states during the period 1999-2013 by considering also inequalities and demographics. He found that the elasticity of consumption out of housing wealth is 0.05 for the whole sample and 0.075 for poor households. For the high loan-to-value mortgages households' elasticity of consumption is 0.19. Moreover, the estimated elasticity of consumption tends to be negative or statistically insignificant for younger households.

The literature to date has ignored the impact of the interest rate on consumption decisions. However, this choice is unwarranted from a theoretical point of view given the importance of the interest rate in affecting savings, and hence, consumption decisions (Romer 2012). In particular, interest rate changes cause both a substitution and an income effect which run in opposite directions, thus creating an uncertain effect on consumption. Jansen (2013) investigates wealth effects on consumption in financial crises for Norway and concludes that real interest rates and wealth need to be included in the long-run relationship. This need is associated with their linking with monetary policy. He provides evidence that changes of the correlations between wealth and the real interest rate comove with changes in the monetary policy system.

The U.S. Treasury rate is the sum of two components, the average of expected future short-term rates, and a term premium (Greenlaw et al. 2018). It is also an important economic indicator. Economic announcements affect prices, trading volume and bid-ask spreads. The bid-ask spreads immediately return to normal levels within 5 to 15 minutes. In an experiment Balduzzi et al. (1996) found that only three out of seventeen announcements affect the bill price. Interestingly, for announcements that have a significant impact on prices, the impact occurs within one minute after the announcement. Economic announcements have less effect on trading volume for the three-month bill, although a change in monetary policy leads to an average trading volume up to nine times higher than at non-announcement times. For the 10-year rate, previous research (Balduzzi et al. 1996) found a strong association between announcements and trading volume. Since then, and by the introduction of the unconventional monetary policy, the 10-year treasury rate became an important determinant of consumer decisions (Swanson 2015; Williams 2013; Wright 2012).

Hence, in this study we use the 10-year treasury constant maturity rate as our fourth variable associated with consumption. Since the bond yields represent interest for borrowing money, the yield curve is of tremendous importance, both from a conceptual and a practical point of view. "From a conceptual perspective, the yield curve determines the value that investors place today on nominal payments at all future dates – a fundamental determinant of almost all asset prices and economic decisions. From a practical perspective, the U.S. Treasury market is one of the largest and most liquid markets in the global financial system. In part because of this liquidity, U.S. Treasuries are extensively used to manage interest rate risk, to hedge other interest rate exposures and to provide a benchmark for the pricing of other assets" (Gürkaynak et al. 2006).

3 Model and data

We use panel data for the 50 US states to obtain a US consumption function for several reasons. First, obviously by combining cross-sectional and time series data, we increase the size of information and the number of degrees of freedom, and thus the accuracy of estimation. We introduce information on the dynamic behaviour of many entities at the same time. Moreover, panel data allow us to use Westerlund cointegration, which considers the cross-sectional dependence and the heterogeneity of the states. Finally, we can remove the impact of omitted variables bias in the regression results and hence obtain more reliable results. Note that the results of the cointegration test stand for the US and not for the US states separately.

Based on the discussion above, consumption is described as a function of income, financial wealth, housing wealth and the 10-year treasury rate. The empirical model takes the following log-linear form where all variables, except for the interest rate, are expressed in logs:

$$\ln \mathrm{C}_{\mathrm{it}} = \beta_1 \ln \mathrm{Y}_{\mathrm{it}} + \beta_2 \ln \mathrm{F\,W}_{\mathrm{it}} + \beta_3 \ln \mathrm{H\,W}_{\mathrm{it}} + \beta_4 \mathrm{MRIN}_{\mathrm{it}} + \mathrm{FEffects} + \mathrm{e}_{\mathrm{it}} \quad (1)$$

where,

$\ln C_{it}$: (log) real per capita consumption in state i at time t,

$\ln Y_{it}$: (log) real per capita personal income in state i at time t,

$lnFW_{it}$: (log) real per capita financial wealth in state i at time t,
$lnHW_{it}$: (log) real per capita housing wealth in state i at time t,
$MRIN_{it}$: 10-year treasury constant maturity rate in US, and
e_{it}: disturbance term.

The data used in this study consists of quarterly, state-level panel observations spanning from the 1st quarter of 1975 to 1st quarter of 2018 for consumption, personal income, housing wealth, financial wealth and 10-year treasury for the 51 U.S. states. The data is constructed in the spirit of Case et al. (2005, 2011, 2012).

Figure 1 plots the US series of consumption, income, financial and housing wealth. Housing wealth follows an upward trend until the decline instigated by the 2008 crisis, while consumption follows mostly the housing wealth trend up to that period and after the huge decline of 2008 it almost remains stable up to now. Housing wealth also exceeded financial wealth during the three recessions experienced in 1980, 1981-1982, and 2007-2009.

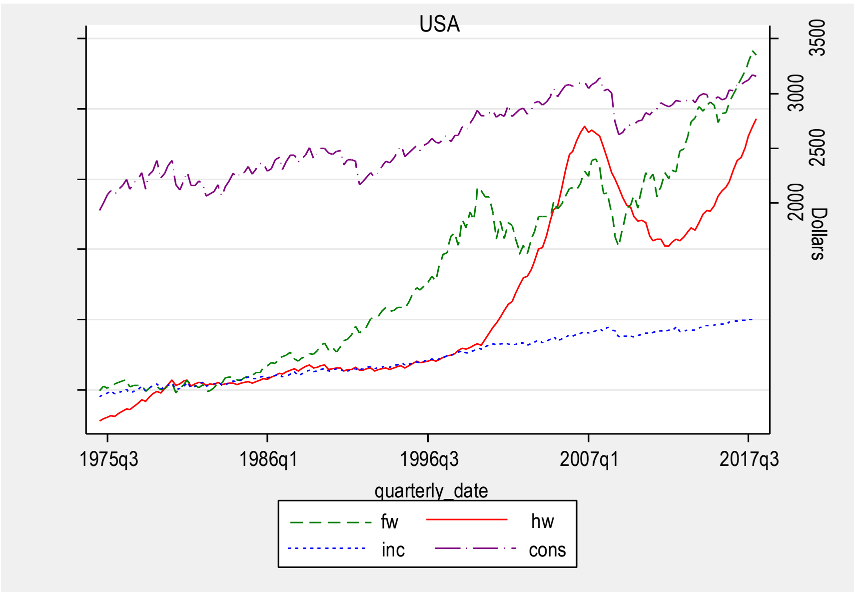

Figure 1. US Consumption, income, financial and housing wealth, during the period 1975-2018, (per capita of US dollars). The right-hand scale measures consumption and the left-hand scale measures income, financial and housing wealth.

All variables are in chained 2005 dollars and measured per capita. Figure 2 presents the 10-year treasury constant maturity rate in daily frequency obtained from the Federal Reserve Bank of St. Louis (FRED).

Figure 2: US 10-Year Treasury Constant Maturity Rate (7/3/1975-7/3/2019 in daily frequency)

On June 1, 2012, the 10-year treasury rate dropped to 1.46 percent which was its lowest level in 200 years. This low treasury interest rate level was attributed to investors' high demand for treasuries associated with the Eurozone debt crisis. Furthermore, on July 5, 2016, the 10-year treasury reached another low record in its history of 1.375 percent arising from the low inflation rate and the UK's referendum vote for Brexit (The CNBC Journal, 2016).

We proceed by presenting our data for consumption. We use retail sales as proxy for consumption. In the spirit of Case, Quigley and Shiller (2011, 2012) who stated that there are no measures of consumption spending by households recorded at the state level, we have chosen retail sales as a proxy of household spending. Retail sales account for roughly half of the total consumer expenditures" (Case et al., 2012, p. 113). We source monthly, seasonally adjusted national advance retail sales (excluding food services) from FRED, as state-level retail sales data is not available. We exclude food services from our data because the change in income or wealth may not adjust the consumption of food services in the same direction. Since retail sales data are not available for US states, we use the allocation of retail trade data for each state to obtain the state-level retail sales. Hence, we follow two steps: First, we calculate the proportion of retail trade of each state in the US total. We then apply these proportions to US retail sales to determine the state retail sales data. Personal income data and state retail trade data are obtained from the Bureau of Economic Analysis (BEA), published on June 21st, 2018.

The series were deseasonalized using the CensusX-12 method where necessary. The housing wealth variable is constructed through the following equation:

$$V_{it} = R_{it} N_{it} I_{it} V_{io} \qquad (2)$$

where,
V_{it} is the aggregate housing wealth of owner-occupied housing in state i in quarter t,
R_{it} is the homeownership rate in state i in quarter t,
N_{it} is the number of households in state i in quarter t,
I_{it} is the weighted repeat sales price index, for state i in quarter t, and finally V_{io} is the mean home price for state i in the base year, 2000.

The mean home price historical data from 1975 to 2005 every decade is obtained from Census and it is transformed to quarterly data through a linear interpolation method based on the previous and next non-missing values. Annual data from 2005 to 2017 are obtained from Factfinder of Census, while for the first quarter of 2018 we use the rises given by Zillow real estate (https://www.zillow.com/wy/home-values/).

The homeownership rate and the data on the number of households are obtained from the Census Bureau and Factfinder of Census Bureau, respectively. We obtain the data of the number of households from 2016 to 2018 through forecasting techniques. For the weighted repeat sales price index, we use the all-transactions house price index for each state obtained by FRED. We convert the base year from 1980 to 2000 to match better our data. Finally, total financial wealth consists of the sum of corporate equity, mutual fund shares and pension funds held by households and is obtained from FRED. The allocation of the financial market (obtained by BEA) is used to allocate data to each state due to data unavailability at state level.

4 Methods and findings

4.1 Panel unit root analysis

The first and crucial step in our analysis is to determine the order of integration of the variables. Typically, the linear combination of I(1) variables will itself be I(1), unless the variables are cointegrated. There is a variety of tests for unit roots or stationarity in panel datasets which are classified into two groups. The first-generation tests assume cross-sectional independence while the second-generation tests take into account the dependence across the different units in the panel (cross-sectional dependence). The cointegration analysis becomes more problematic when the unit roots in the different cross section units are due to common random walk components (Breitung and Pesaran 2005). Ignoring cross-sectional dependence of errors can have serious consequences, and the presence of some form of cross-sectional correlation of errors in panel data applications in economics is likely to be the rule rather than the exception (Baltagi, 2015).

Therefore, we first test for the presence of panel cross-sectional dependence using the tests proposed by Frees (1995), Friedman (1937) and Pesaran (2006). Table 1 presents the results of the three tests for cross-sectional independence. All tests strongly reject the null hypothesis of cross-sectional independence in our panel.

Having established evidence for cross-sectional dependence, we apply Pesaran's (2007) CIPS unit root test under cross-sectional dependence. Table 2 reports strong evidence that all variables are integrated of order one in two specifications: (a) a model that includes a constant term and (b) a model that includes both a constant and a trend. Hence, the next step is to investigate the possibility of a long-run or cointegrating relation among consumption, personal income, financial wealth, housing wealth, and the 10-year treasury rate.

4.2 Panel cointegration analysis

We conduct two types of cointegration tests to test for the presence of a long-run relationship among our integrated variables. First, we use Pedroni's (1999, 2004) approach which proposes a residual-based cointegration test for cross-sectionally independent panel and, second, we employ Westerlund's (2007) panel cointegration test which allows for cross-sectional dependence.

Pedroni (1999) developed seven cointegration statistics which allow for considerable heterogeneity among individual members of the panel, in both the long-run cointegrating vectors and the dynamics associated with short-run deviations from these cointegrating vectors. Dynamic OLS involves adding lags and leads of the regressors to eliminate feedback effects and endogeneity. Additionally, these tests are appropriate both for the case with common autoregressive roots and for heterogeneity of the autoregressive root under the alternative hypothesis in the spirit of Im, Pesaran, and Shin (1997). The most significant reason cited in empirical literature for using Pedroni test is the increased power arising from accounting both the time series and cross-sectional dimensions. Despite this, many researchers, such as Ho (2002), failed to reject the null hypothesis of no cointegration and supported that the international capital mobility is high although theory suggested the opposite. The explanation of this failure lies on the common factor restriction, which requires that the long-run cointegrating vector for the variables in their levels must be equal to the short run adjustment process for the variables in their differences.

Results for the panel cointegration tests are presented in Table 3. All seven test statistics are distributed as N(0,1), under the null of no cointegration. The tests were performed for two models: a model including a constant and a model including both a constant and a trend. Almost all test statistics reject the null of no cointegration hypothesis at one percent significance level. This implies that consumption, financial wealth, housing wealth and the 10-year treasury rate are linked by a long-run equilibrium relationship.

In the next step, we present the results of the long-run equilibrium relationships using the estimation method of Dynamic OLS (DOLS) for heterogeneous panels (Pedroni, 2004). Results of the group mean average regression are shown in Table 4. Two lags and leads are selected by the Akaike information criterion. The estimated long-run coefficients are all statistically significant except for the long-run interest rate. As expected, income and the two wealth proxies have a positive effect on consumption. The estimated income elasticity is 0.87 and it is much larger than the wealth elasticities. According to our results, financial and housing wealth have a similar impact on consumption.

Given our evidence for cross-sectional dependence, the results of the Pedroni cointegration tests are not highly informative. Westerlund (2007) proposed four new panel tests based on structural dynamics, to test the null of no cointegration in the presence of cross-sectional dependence. These tests do not impose any common factor restriction. They are designed to test whether the error correction term in a conditional error-correction model is equal to zero. If this term is statistically different from zero, then the null hypothesis of no cointegration is rejected. Each test involves individual specific short-run dynamics with serially correlated error terms, non-strictly exogenous regressors, individual-specific intercept and trend terms, individual specific slope parameters and cross-sectional dependence. Bootstrap tests are also proposed to handle applications with cross-sectional dependence. The error-correction tests assume the following data-generating process:

$$\Delta y_{it} = \delta' d_t + \lambda_i\left(y_{i,t-1} - \beta_i' x_{i,t-1}\right) + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \gamma_{ij} \Delta x_{i,t-j} + e_{it} \tag{3}$$

where t=1,. . . ,T and i=1,. . . ,N are the time and the sectional units, respectively, p_i and q_i are the lags and leads[1] orders and are permitted to vary across individuals,

[1] 'By adding leads and not just lags of Δxit, we can allow for regressors that are weakly but not necessarily strictly exogenous' (Westerlund 2007, p. 234).

allowing for a heterogeneous serial correlation structure. The parameter λ_i determines the speed at which the system returns to the equilibrium after a sudden shock. If $\lambda_i < 0$ and statistically significant, then the variables are cointegrated. If $\lambda_i = 0$, there is no cointegration. We assume that errors are independent across both i and t. We handle any dependence across i by means of bootstrap methods (Westerlund 2007).

Based on the general form of Equation (3), we specify the estimated error-correction equation for consumption as follows:

$$\begin{aligned} \Delta C_{it} = a_i^C + \lambda_i^C C_{i,t-1} + (-\lambda_i^C)\beta_{1i}^c Y_{i,t-1} + (-\lambda_i^C)\beta_{2i}^c FW_{i,t-1} + (-\lambda_i^C)\beta_{3i}^c HW_{i,t-1} \\ + (-\lambda_i^C)\beta_{4i}^c MR_{i,t-1} \\ + \sum_{j=0}^{p_i} \theta_{i,j}^c \Delta Y_{i,t-j} + \sum_{j=1}^{p_i} \varphi_{i,j}^c \Delta C_{i,t-j} + \sum_{j=0}^{p_i} \delta_{i,j}^c \Delta FW_{i,t-j} \\ + \sum_{j=0}^{p_i} \zeta_{i,j}^c \Delta HW_{i,t-j} + \sum_{j=0}^{p_i} \eta_{i,j}^c \Delta MR_{i,t-j} + e_{it} \end{aligned} \tag{4}$$

Here, the parameter vector λ^v; $v \in (C; Y; FW; HW; MR)$ refers to the set of error-correction equations and includes the error-correction term parameters. It provides estimates of the speed of error correction towards the long-run equilibrium for each state. The short-run coefficients are the parameters that refer to the first-differenced lagged variables. For example, the short-run income effects on consumption are captured by the θ's.

Table 5 presents the results of the four Westerlund test statistics for the no cointegration null and estimates of the cointegration (or long run) coefficients. This table also reports the short-run coefficients and estimates of the error-correction terms. The tests are carried out with 100 bootstrap replications. The four test statistics differ according to the criterion used in measuring the number of lags and leads and the bandwidth. The number of lags/leads is chosen according to two criteria: First, the Akaike information criterion, and second, the T rule[2] as a function of T. Additionally, all tests are constructed with bandwidth chosen according to the rule recommended by Newey and West (1994). Several interesting conclusions are obtained from the results reported in Table 5. First, there is strong evidence for cointegration under most test statistics and in most cases of the lag specification. Second, the coefficient of the error-correction term is negative and significant at 1% in all four estimated models and ranges

[2] It holds under the function: $2(T/100)^{2/9}$

from 0.190 to 0.281 in absolute value. The results depend on the number of the imposed lags, leads and bandwidth. Third, the estimated income elasticity is positive and significant and varies between 0.62 and 1.08, depending on the model specification. Fourth, the estimated housing wealth elasticity is positive and significant at 1% in all specifications with size varying between 0.07 and 0.12. Fifth, the estimated financial wealth elasticity is in general insignificant except for one case where the sign is negative and significant at 5%. Finally, the effect of the long-term interest rate on consumption is negative and insignificant except for one case where significance obtains. The interpretation of the interest rate coefficient suggests that an increase of one percentage unit in the treasury rate decreases consumption by 0.32%.

It would be interesting to compare the estimated income and wealth elasticities across the different estimation methodologies. Estimates of the income elasticity of consumption are 0.87, under both the Pedroni DOLS estimation and Westerlund cointegration test, when lags are chosen according to the AIC. According to the Westerlund cointegration test, the elasticity of consumption with respect to income is on average 0.92. In Pedroni's DOLS, the estimates of the elasticity of consumption for housing wealth and financial wealth are relatively similar (0.059 vs. 0.044). Under the Westerlund methodology, the estimated elasticity of consumption for housing wealth is 0.07 in most specifications, with the three lags specification giving the highest elasticity of 0.115. In contrast, financial wealth seems to have an insignificant impact on consumption. Finally, the long-term interest rate has a statistically insignificant effect on consumption under most cases in the Westerlund estimations. Similarly, the effect is insignificant under the Pedroni estimation.

In Table 5 we also report the short-run effects of each determinant on consumption. Short-run effects are captured by the coefficients of the variables in lagged first differences. In all the estimated models, the short-run effects of income, financial and housing wealth on consumption are positive and statistically significant at 1%. As in the case of the long-run effects, income in the short run has a larger impact on consumption than both measures of wealth, financial and housing wealth. Both financial and housing wealth have a similar positive effect on consumption. However, for financial wealth there is an important difference between short and long run. In the short run, as opposed to the long run, financial wealth generates a strong positive effect on consumption. As in the case of the long-run effects, the evidence suggests that housing wealth has a bigger effect on consumption than that of financial wealth. Across all estimated specifications, a change of 1% in financial wealth causes a change of 0.059% in consumption in the same direction. This result is slightly below that of housing wealth. In the short run, a one percentage unit change of housing wealth generates increases in consumption between 0.062% and 0.092% with an average of 0.072%. In marked contrast, the short-term financial wealth effect appears stronger than that of the long run and eventually fades out.

4.3 Panel causality analysis

Given the evidence for cointegration, we perform causality analysis in order to investigate causal interactions among the variables. We perform two causality tests. First, we test for causality by performing the Westerlund cointegration test, thus regressing each variable against the other four to obtain the causal linkages in our panel data. The panel VECM can be written as follows:

$$\begin{aligned} \Delta Y_{it} = {} & a_i^Y + \lambda_i^Y Y_{i,t-1} + (-\lambda_i^Y)\beta_{1i}^Y C_{i,t-1} + (-\lambda_i^Y)\beta_{2i}^Y FW_{i,t-1} + (-\lambda_i^Y)\beta_{3i}^Y HW_{i,t-1} \\ & + (-\lambda_i^Y)\beta_{4i}^Y MR_{i,t-1} \\ & + \sum_{j=1}^{p_i} \theta_{i,j}^Y \Delta Y_{i,t-j} + \sum_{j=0}^{p_i} \varphi_{i,j}^Y \Delta C_{i,t-j} + \sum_{j=0}^{p_i} \delta_{i,j}^Y \Delta FW_{i,t-j} \\ & + \sum_{j=0}^{p_i} \zeta_{i,j}^Y \Delta HW_{i,t-j} + \sum_{j=0}^{p_i} \eta_{i,j}^Y \Delta MR_{i,t-j} + e_{it} \end{aligned} \tag{5}$$

$$\begin{aligned} \Delta FW_{it} = {} & a_i^{FW} + \lambda_i^{FW} FW_{i,t-1} + (-\lambda_i^{FW})\beta_{1i}^c Y_{i,t-1} + (-\lambda_i^{FW})\beta_{2i}^{FW} C_{i,t-1} \\ & + (-\lambda_i^{FW})\beta_{3i}^{FW} HW_{i,t-1} + (-\lambda_i^{FW})\beta_{4i}^{FW} R_{i,t-1} \\ & + \sum_{j=1}^{p_i} \theta_{i,j}^{FW} \Delta FW_{i,t-j} + \sum_{j=0}^{p_i} \varphi_{i,j}^{FW} \Delta Y_{i,t-j} + \sum_{j=0}^{p_i} \delta_{i,j}^{FW} \Delta C_{i,t-j} \\ & + \sum_{j=0}^{p_i} \zeta_{i,j}^{FW} \Delta HW_{i,t-j} + \sum_{j=0}^{p_i} \eta_{i,j}^{FW} \Delta R_{i,t-j} + u_{it} \end{aligned} \tag{6}$$

$$\begin{aligned} \Delta HW_{it} = {} & a_i^{HW} + \lambda_i^{HW} HW_{i,t-1} + (-\lambda_i^{HW})\beta_{1i}^c C_{i,t-1} + (-\lambda_i^{HW})\beta_{2i}^c Y_{i,t-1} + (-\lambda_i^{HW})\beta_{3i}^{HW} FW_{i,t-} \\ & + (-\lambda_i^{HW})\beta_{4i}^{HW} R_{i,t-1} \end{aligned} \tag{7}$$

$$+\sum_{j=1}^{p_i}\theta_{i,j}^{HW}\,\Delta HW_{i,t-j}+\sum_{j=0}^{p_i}\varphi_{i,j}^{HW}\,\Delta C_{i,t-j}+\sum_{j=0}^{p_i}\delta_{i,j}^{HW}\Delta Y_{i,t-j}$$
$$+\sum_{j=0}^{p_i}\zeta_{i,j}^{HW}\Delta FW_{i,t-j}+\sum_{j=0}^{p_i}\eta_{i,j}^{HW}\Delta R_{i,t-j}+\varepsilon_{it}$$

$$\Delta R_{it}=a_i^R+\lambda_i^R R_{i,t-1}+(-\lambda_i^R)\beta_{1i}^R C_{i,t-1}+(-\lambda_i^R)\beta_{2i}^c Y_{i,t-1}+(-\lambda_i^R)\beta_{3i}^R FW_{i,t-1}+(-\lambda_i^R)\beta_{4i}^R HW_{i,t-1}+\sum_{j=1}^{p_i}\theta_{i,j}^{R} R_{i,t-j}+\sum_{j=0}^{p_i}\varphi_{i,j}^{R}\Delta\, C_{i,t-j}+\sum_{j=0}^{p_i}\delta_{i,j}^{R}\Delta\, Y_{i,t-j}+\sum_{j=0}^{p_i}\zeta_{i,j}^{R}\Delta FW_{i,t-j}+\sum_{j=0}^{p_i}\eta_{i,j}^{R}\Delta HW_{i,t-j}+z_{it} \quad (8)$$

Based on the above specifications, we can test for both short-run and long-run causality. We examine the evidence for short-run causality by testing through the Wald restriction test the joint significance of the parameters for the variables in their lagged differences. We examine the evidence for long run causality as implied by the statistical significance of the t-statistics on the error-correction parameter.

Additionally, we also employ the panel causality methodology suggested by Dumitrescu & Hurlin (2012), thereafter DH. Individual Wald statistics of Granger noncausality converge sequentially to a standard normal distribution characterized for a fixed T sample. Monte Carlo experiments show also that these standardized panel statistics have very good small sample properties, even in the presence of cross-sectional dependence (DH 2012). DH, based on the seminal paper of Granger (1969), provide an extended test designed to detect causality in panel data.

Table 6 presents the results of the Granger causality based on the Westerlund cointegration model under the four alternative lag specifications reported in Table 5. Table 7 summarizes the corresponding directions among the variables. In the first estimated model where the variables are estimated with 3 lags, personal income exhibits bidirectional causality with consumption at 1% significance level, and financial wealth Granger causes consumption at the 5% level, whereas housing wealth and the 10-year treasury rate are insignificant. According to equation (5), the 10-year treasury rate Granger causes personal income at 10% significance level. In two equations personal income and financial wealth exhibit bidirectional causality at 5% significance level. Further, in equation (7) consumption and income Granger cause housing wealth at 10 and 5 %, respectively.

Moreover, the error correction term is statistically significant at the 1% level in all regressions with a fast speed of adjustment to long-run equilibrium in the first three regressions, where consumption, income and financial wealth are the dependent variables (equations 4-6). The speed of adjustment toward long-run equilibrium appears much faster in equation (5) than in equations (4) and (6). Equally, the estimated coefficients in the specifications with 4, 5 and 6 lags indicate similar results.

Table 8 reports the results of Granger causality tests under the DH (2012) methodology. These results are summarized in Table 9. When the optimal number of lags are chosen by Bayesian Information Criterion (BIC), two bidirectional causal relationships exist: between income and consumption, and income and housing wealth. Moreover, financial wealth Granger causes income and housing wealth. Each of the Treasury rate and consumption also Granger causes housing wealth.

The optimal number of lags chosen according to BIC is different for each equation, while in the first equation it is even different for each variable. Consequently, we specify an additional model with four lags for all equations and variables. Table 9 also includes the results by imposing four lags in all variables (4). The results indicate more bidirectional linkages among the variables.

5. Discussion of the results

Our findings appear to be in line with the results of most empirical studies in the literature, using alternative econometric methods and data, such as those by Carrol (2003) and Carrol et al., (2011, 2017), Rapach and Strauss, (2006), Mian (2013), Case et al. (2012) and Bampinas et al. (2017). We provide evidence for positive long-run effects of income and housing wealth on consumption and short-run effects of financial wealth that stand stronger than those in the long run. The volatility in financial wealth may cause an immediate large effect on consumption which eventually fades out. Consumption exhibits larger response to housing wealth changes compared to those of financial wealth. The elasticity of consumption with respect to housing wealth is getting stronger over time compared to that of financial wealth which seems to be negligible. Our findings confirm that housing wealth effects on consumption in the short run are significant and large and the coefficient ranges from 0.062 to 0.092. In the long run, the estimated elasticity is between 0.072 and 0.154 indicating the persistence of housing wealth over time and its high significance in consumption growth.

According to Slacalek (2009), the US, among other countries, such as Belgium and the Netherlands, possess more financial wealth than housing wealth. Consequently, the aggregate effect of financial wealth on consumption appears larger in the short run simply because these countries have more financial wealth. In line with this background, our data shows that the periods in which housing wealth surpassed financial wealth were 1980-1985 and the recent Great Recession. Gabriel et al. (2008) suggest that short-term deviations in the consumption-wealth ratio will forecast either asset returns or consumption growth; the first when changes in wealth are transitory, the second when changes in wealth are permanent.

As a result, consumption responses to financial wealth shocks appear stronger in the short run than the elasticity of consumption for housing wealth and gradually fade out. Figure 1 documents the finding of most literature (Helbling and Terrones 2003; Slacalek 2009) that housing wealth follows a smoother path than financial wealth. "In contrast to stock prices, when housing prices fall, they typically do so gradually over several quarters or even years rather than days" (Slacalek 2009, p. 15).

Equally, our findings report the elasticity of consumption with respect to housing wealth to be greater than that of the stock market wealth in accordance with most of the literature. It may be worth mentioning that the short-run coefficient of income starts from 0.640 and rises over time getting to an average of 0.865 for all times according to the Pedroni test and to 0.870 according to the Westerlund test, a finding in line with Kuznets (1946). In his seminal work, he found that the proportion of consumption to income was equal to about 0.87 in three sub-samples during the period 1869-1933.

With respect to the 10-year treasury rate, the estimated coefficient appears positive and relatively large in the short run, whereas it is getting negative or/and insignificant in the long run. We find that for a one-unit increase in the 10-year treasury rate, consumption rises by about 40 cents in the short run. Our evidence for a positive short-run association between the interest rate and consumption is in line with standard textbook analysis that emphasises the interplay of the substitution and income effect of an interest rate change (Romer 2012). Moreover, this result supports consumer behavior influenced by expectations of future economic policies that affect consumer decisions only in the short run. This second explanation is more in line with our evidence that the interest rate is insignificant in the long run[3]. In addition, this result squares with the literature (Swanson 2015; Williams 2013; Wright 2012). Note the

[3] To test whether this interpretation is consistent with our empirical analysis, we use trade volume data which reflect policy news and test for a short-run and a long-run relationship with the 10-year Treasury rate. We find that the interest rate Granger-causes trade volume but there is no long-run relationship between the two variables.

strong association between the 10-year treasury rate and trading volatility which is an indicator for information flow (Balduzzi et al. 1996).

In recent years, the 10-year treasury rate is associated with large scale asset purchase announcements and provides information about interest rate expectations and risk premia over long horizons of 10 years (Swanson 2015). Besides, long-term interest rates are – among other macro-indicators, such as, long-run inflation and unemployment rates – crucial in explaining the Federal Funds Rate (Wölfel & Weber 2017). Under the unconventional policy, the monetary authority compresses the spread (e.g. the difference between the long-term bond and short-term bill rate) via large-scale asset purchases to support economic activity and promote higher inflation. However, large-scale asset purchases by the central bank seem to signal further purchases in the future. Asset prices do not significantly impact unconventional monetary policy measures (Agnello, Castro, Dufrénot, Jawadi, and Sousa 2018).

Regarding the Granger causality tests, the strongest bidirectional causality exists among consumption and income. Generally, our findings support the short-run adjustments of cointegration tests. Moreover, the short-run causality analysis implies that housing wealth has a predictive power in forecasting the 10-year treasury rate, or equivalently, the dynamics of housing wealth provide information on the future movements of interest rates.

Consequently, there is considerable evidence of the importance of the housing wealth channel in consumption movements. Wealth may reflect a new view of future profits in an increase in the treasury rate. Hence, this study confirms previous studies (Mishkin 2007) that interest rates directly influence expectations of future house price movements, and housing supply; and indirectly influence the real economy.

6 Conclusion

This paper examines the empirical relationship among consumption and its determinants in a panel set up. Using an original data set for the 50 US states and the cointegration methodology of Pedroni (1997) and Westerlund (2007), we investigate the short run and long-run relationship among consumption, income, financial, and housing wealth, and a long-term interest rate. Using both cointegration and error-correction analysis we obtain the following results. First, in the long run, we find strong evidence for positive income and housing wealth effects on consumption. The estimated income elasticity varies between 0.62 and 1.08, depending on the model specification and the estimated housing wealth elasticity varies between 0.07 and 0.12.

Financial wealth and the long-term interest rate do not seem to affect consumption significantly in the long run. Second, in contrast to the long-run effects, in the short run, both forms of wealth and the interest rate are also significant and have a positive and negative effect on consumption, respectively. However, the effect of housing wealth seems to be larger than that of financial wealth.

We also test for short-run and long-run causality between consumption and its determinants. The causality tests indicate a strong bidirectional short-term causality between per capita consumption, income and financial wealth and a long-term causality between all the variables. The long-run causality analysis also shows that sudden shocks arising from variables other than income, i.e., financial wealth, consumption and the 10-year treasury, have an immediate effect on housing but the speed at which will eventually be offset is slower compared to the other systems where the dependent variable is consumption, financial wealth, income or 10-year treasury. Finally, the high significance of housing wealth and its linkage with the 10-year treasury rate in most specifications, confirms that developments in the housing market have a major and predictable effect on real economic activity.

Compliance with Ethical Standards:

Ethical approval: This article does not contain any studies with human participants or animals performed by any of the authors.

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Appendix 1

Table 1. Cross-sectional independence tests

	Cd-test	Frees-test (Q stat.)	Friedman's test
Statistic	135.609	10.931	2545.962
P-Value	0.000	0.000	0.000
Avg of absolute value	0.414	0.414	0.414

Note: The tests boil down to verifying whether sum of correlations between panel units is equal to zero.

Table 2. CIPS Panel Unit Root test

	Pesaran's CADF t-bar statistics			
	4 lags		3 lags	
Variables	Constant	Constant trend	Constant	Constant trend
lnConsumption	-1.921 [0.150]	-2.025 [0.998]	-1.553 [0.970]	-1.681 [1.000]
lnIncome	-2.207 [0.000]	-1.929 [1.000]	-2.233 [0.000]	-1.989 [0.999]
lnFinancialWealth	-1.872 [0.259]	-2.070 [0.994]	-1.797 [0.478]	-2.020 [0.998]
lnHousingWealth	-2.124 [0.004]	-1.908 [1.000]	-1.806 [0.449]	-1.551 [1.000]
ΔlnConsumption	-4.814 [0.000]	-4.897 [0.000]	-5.284 [0.000]	-5.378 [0.000]
ΔlnIncome	-5.277 [0.000]	-5.393 [0.000]	-5.584 [0.000]	-5.707 [0.000]
ΔlnFinancialWealth	-5.557 [0.000]	-5.699 [0.000]	-5.908 [0.000]	-6.419 [0.000]
ΔlnHousingWealth	-3.605 [0.000]	-3.679 [0.000]	-4.328 [0.000]	-4.409 [0.000]

Note: Δ is the first difference operator. Numbers in brackets are p-values. The Akaike Information Criterion was used to determine the optimal lag length. The CIPS test assumes cross-section dependence in the form of a single unobserved common factor.

Table 3. Pedroni Cointegration test statistics

Test	Constant	Constant and Trend
Panel v-statistic	-0.885	0.0872
Panel ρ statistic	0.888	-1.76**
Group PP-statistic t	0.392	-2.78***
Panel ADF-statistic	4.23***	2.873***
Group-statistic	-0.3261	-2.108***
Group PP-statistic (non-parametric)	-0.591	-3.028***
Group ADF-statistic (non-parametric)	4.173***	2.861***

Note: *** indicates statistical significance at 1%.

Table 4. Panel cointegration coefficients (Pedroni's Panel DOLS)

Variables	**Coefficient**	**t-stat**
Ln Income	0.874	32.78***
lnFinancial Wealth	0.0439	1.807***
lnHousing Wealth	0.0589	9.92***
10-Year Treasury Rate	-0.1868	-0.2832

Note: Pedroni's PDOLS (Group mean average). Number of Panel units: 51. Lags and leads: 2 Number of obs.: 8568. Avg obs. per unit: 168. Data has been time-demeaned.

Table 5. Statistics and the panel cointegration coefficients in the Westerlund model

Independent Variables	Long-run coefficient	Short-run coefficient	ECT	Statistics	P-value	Robust P-value	Lags/Leads/ Irwindow
Income	0.624 [0.000] ***	0.790 [0.000] ***	-0.190 [0.000] ***	Gt	0.000	0.020	3/2/4 (T=173)
Financial Wealth	0.1901 [0.138]	0.0813 [0.000] ***		Ga	0.004	0.080	
Housing Wealth	0.1154 [0.009] ***	0.0923 [0.000] ***		Pt	0.072	0.180	
10-Year Treasury Rate	0.822 [0.332]	0.4169 [0.000] ***		Pa	0.058	0.170	
Constant	-1.220 [0.226]	-0.4133[0.001] ***					
Trend	-0.004 [0.018] **	-0.0005[0.000] ***					
Income	0.8708 [0.000] ***	0.6377[0.000] ***	-0.2151 [0.000] ***	Gt	0.000	0.000	4/3/4 (AIC)
Financial Wealth	-0.0252[0.449]	0.0604[0.000] ***		Ga	0.002	0.150	
Housing Wealth	0.0720[0.000] ***	0.0718[0.000] ***		Pt	0.000	0.030	
10-Year Treasury Rate	-0.1400[0.153]	0.4919[0.000] ***		Pa	0.004	0.140	
Constant	-1.2530[0.113]	-0.3967[0.003] ***					
Trend	-0.0017[0.000] ***	-0.0004[0.000] ***					
Income	1.096[0.000] ***	0.640[0.000] ***	-0.254 [0.000] ***	Gt	0.000	0.000	5/4/10 (T=8823)
Financial Wealth	-0.090[0.044] **	0.051[0.000] ***		Ga	0.656	0.360	
Housing Wealth	0.071[0.000] ***	0.062[0.000] ***		Pt	0.001	0.020	
10-Year Treasury Rate	-0.329[0.001] ***	0.413[0.000] ***		Pa	0.697	0.460	
Constant	-2.823[0.004] ***	-0.898[0.000] ***					
Trend	-0.002[0.000] ***	-0.001[0.000] ***					
Income	1.084[0.000] ***	0.685[0.000] ***	-0.281 [0.000] ***	Gt	0.000	0.000	6/5/11 (T=8823)
Financial Wealth	-0.137[0.124]	0.044[0.000] ***		Ga	0.904	0.400	
Housing Wealth	0.076[0.000] ***	0.067[0.000] ***		Pt	0.161	0.060	
10-Year Treasury Rate	-0.226[0.217]	0.445[0.000] ***		Pa	0.965	0.540	
Constant	-2.298[0.028] **	-0.923[0.000] ***					
Trend	-0.001[0.108]	-0.001[0.000] ***					

Note: P-values are in brackets and ***, ** are estimated value significant at the 1\% and 5\% level respectively. The tests are carried out with 100 bootstrap replications. The number of lags is chosen according to AIC and the T rule.

Table 6. Panel causality test results based on the Westerlund cointegration model

Dependent Variables	**Sources of causation (Independent Variables)**											
	SHORT RUN										LONG RUN	
	ΔConsumption		ΔIncome		ΔFinancial Wealth		ΔHousing Wealth		Δ10-Year Treasury Rate		ECT	
	3 lags	4 lags	3 lags	4 lags	3 lags	4 lags	3 lags	4 lags	3 lags	4 lags	3 lags	4 lags
ΔConsumption [Eq. (4)]	–	–	21.87 (-) [0.0000]	23.04 (+) [0.0000]	3.59 (+) [0.0155]	3.95 (+) [0.0047]	0.22 (+) [0.8814]	0.38 (+) [0,8234]	0.74 (-) [0.5308]	1.67 (-) [0.1613]	-0.190 [0.000]	-0.215 [0.0000]
ΔIncome [Eq. (5)]	6.77 (-) [0.0003]	7.64 (-) [0.0000]	–	–	2.67 (-) [0.0500]	3.37 (+) 0.0117]	3.72 (-) [0.0132]	2.66 (-) [0.0356]	2.44 (-) [0.0675]	3.57 (-) [0.0087]	-0.235 [0.000]	-0.218 [0.0000]
ΔFinancial Wealth [Eq. (6)]	2.06 (-) [0.1087]	2.65 (-) [0.0365]	8.94 (-) [0.0001]	8.37 (-) [0.0000]	–	–	0.41 (+) [0.7452]	0.43 (+) [0.7858]	0.34 (-) [0.7931]	1.05 (-) [0.3837]	-0.143 [0.0000]	-0.166 [0.0000]
ΔHousing Wealth [Eq. (7)]	2.23 (-) [0.0872]	0.94 (-) [0.4405]	6.60 (-) [0.0002]	2.095 (+) [0.210]	0.78 (-) [0.5051]	0.82 (-) [0.5147]	–	–	1.10 (-) [0.3522]	0.86 (-) [0.4925]	-0.047 [0.0000]	-0.048 [0.0000]
Δ10-Year Treasury Rate [Eq. (8)]	0.26 (-) [0.8556]	0.37 (-) [0.8296]	0.02 (+) [0.9970]	0.07 (-) [0.9919]	0.26 (+) [0.8523]	0.19 (-) [0.9412]	2.92 (+) [0.0365]	2.30 (+) [0.0622]	–	–	-0.099 [0.0000]	-0.078 [0.0000]
	5 lags	6 lags	5 lags	6 lags	5 lags	6 lags	5 lags	6 lags	5 lags	6 lags	5 lags	6 lags
ΔConsumption [Eq. (4)]	–	–	26.72 (+) [0.000]	24.72 (+) [0.0000]	4.99 (+) [0.0004]	5.64 (+) [0.0000]	0.10 (+) [0.9914]	0.32 (+) [0.9256]	1.33 (-) [0.2555]	0.97 (-) [0.4522]	-0.254 [0.0000]	-0.281 [0.0000]
ΔIncome [Eq. (5)]	11.01 (-) [0.0000]	12.16 (-) [0.0000]	–	–	4.12 (+) [0.0018]	4.85(+) [0.0002]	0.82 (-) [0.5376]	0.74 (-) [0.6150]	2.99 (-) [0.0142]	2.12 (-) [0.0569]	-0.229 [0.0000]	-0.255 [0.0000]
ΔFinancial Wealth [Eq. (6)]	4.03 (-) [0.0021]	5.17 (-) [0.0001]	8.19 (-) [0.0000]	8.31 (-) [0.0000]	–	–	0.13 (+) [0.9859]	0.36 (+) [0.9000]	1.74 (-) [0.1316]	2.03 (-) [0.0686]	-0.203 [0.0000]	-0.224 [0.0000]
ΔHousing Wealth [Eq. (7)]	0.70 (-) [0.6254]	0.63 (-) [0.7043	7.51 (-) [0.0000]	7.35 (-) [0.0000]	1.12 (-) [0.3545]	1.55 (-) [0.1685]	–	–	1.10 (-) [0.3633]	1.33 (+) [0.2529]	-0.056 [0.0000]	-0.060 [0.0000]
Δ10-Year Treasury Rate [Eq. (8)]	0.06 (-) [0.9978]	0.13 (-) [0.9915]	0.10 (-) [0.9923]	0.22 (-) [0.9690]	0.13 (+) [0.9852]	0.49 (+) [0.8160]	2.60 (-) [0.0292]	3.04 (+) [0.0090]	–	–	-0.081 [0.0000]	-0.096 [0.0000]

Notes: F-statistics reported with respect to short-run changes in elasticity in the independent variables. Probability values are in brackets and reported underneath the F-statistic. ECT represents the coefficient of the error correction term with the P-values in brackets. For the co-joint test, we used the Wald-test. Positive or negative signs of the causality are reported in the parenthesis

Table 7. Direction of Causality

Methodology Westerlund (2007) Cointegration Model / Wald Test

Number of lags: 3	Number of lags: 4	Number of lags: 5	Number of lags: 6
Housing Wealth → 10 Year Treasury Rate **10 Year Treasury → Rate Income** **Income ↔ Financial Wealth** **Consumption ↔ Income**			
Income → Housing Wealth	Housing Wealth → Income	Income → Housing Wealth	
	Financial ↔ Wealth Consumption		
			10 Year Treasury Rate → Financial Wealth

Table 8. Dumitrescu & Hurlin (2012) Granger non-causality test results

Optimal number of lags selected by BIC / lags tested: 1 to 55

Dependent Variables	Sources of causation (Independent Variables)				
	ΔConsumption	ΔIncome	ΔFinancial Wealth	ΔHousing Wealth	Δ10Year TreasuryRate
ΔConsumption	–	77.214 {4} (74.684) [0.000]	-3.492 {1} (-3.479) [0.160] [0.150]	2.1808 {1} (2.076) [0.240]	0.558 {1} (0.487) [0.910] [0.920]
ΔIncome	30.285 {4} (29.217) [0.000]	–	29.141 {4} (28.109) [0.000]	16.761{4} (16.115) [0.000]	-2.183 {4} (-2.239) [0.600]
ΔFinancial Wealth	-2.419 {1} (-2.429) [0.420]	4.003 {1} (43.860) [0.130] [0.140]	–	-2.333 {1} (-2.344) [0.150] [0.140]	–
ΔHousing Wealth	34.505 {4} (33.306) [0.000]	65.646 {4} (63.477) [0.000]	6.892 {4} (6.553) [0.000]	–	6.230 {4} (5.917) [0.030]
Δ10-Year TreasuryRate	4.054 {1} (3.911) [0.350] [0.390]	-3.535 {1} (-3.521) [0.510]	-3.462 {1} (-3.450) [0.390]	-0.103 {1} (-0.160) [1.000] [0.930]	–

Notes: Z-statistics reported with respect to four period changes in the independent variables. Z-tilde-statistics are denoted in parentheses. Probability values are in square brackets and reported underneath the partial z-statistic and z-tilde-statistic in order. Single bracket means the same p-value for both statistics. The number of lags chosen by BIC reported in curly brackets. P-values computed using 100 bootstrap replications.

Optimal number of lags: 4

ΔConsumption	–	77.214 (74.684) [0.000]	7.9759 (7.603) [0.050]	17.822 (17.143) [0.000]	2.231 (2.037) [0.730]
ΔIncome	30.285 (29.216) [0.000]	–	29.141 (28.109) [0.000]	16.761 (16.115) [0.000]	-2.183 (-2.239) [0.570] [0.560]
ΔFinancial Wealth	49.623 (47.959) [0.000]	64.867 (62.721) [0.000]	–	15.916 (15.295) [0.000]	-5.177 (-5.140) [0.240] [0.230]
ΔHousing Wealth	34.505 (33.306) [0.000]	65.646 (63.476) [0.000]	6.8923 (6.553) [0.020]	–	6.230 (5.912) [0.010]
Δ10-Year TreasuryRate	3.424 (3.193) [0.470] [0.490]	-1.497 (-1.575) [0.820]	-5.177 (-5.140) [0.350] [0.340]	3.447 (3.215) [0.300] [0.310]	–

Notes: Z-statistics reported with respect to four period changes in the independent variables. Z-tilde-statistics are denoted in parentheses. Probability values are in brackets and reported underneath the corresponding partial z-statistic and z-tilde-statistic, respectively. Single bracket means the same value for both statistics. P-values computed using 100 bootstrap replications.

Table 9. Direction of Causality

Methodology: Dumitrescu & Hurlin (2012)

Number of lags: Bayesian Information Criteria	Number of lags: 4
Consumption ↔ Income Income Housing ↔ Wealth 10 Year Treasury Rate → Housing Wealth	
Financial Wealth → Income	Financial Wealth ↔ Income
Financial Wealth → Housing Wealth	Financial Wealth ↔ Income
Consumption → Housing Wealth	Consumption ↔ Housing Wealth
	Financial Wealth ↔ Consumption