Stock Market Wealth Effects in Emerging Economies of Eastern Europe: Evidence from Russia and Ukraine

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Abstract

The article estimates the stock market wealth effects in Russia and Ukraine and attempts to determine whether institutional differences between Russia and Ukraine and both non-EU and EU Eastern European countries can explain differences in the significance and magnitudes of their wealth effects. Employing the VECM, I find that in the long run, in Russia, the stock market wealth effect is statistically significant at the 5% level, while in Ukraine, it is statistically insignificant. In particular, it is estimated that a 10% increase in stock market wealth increases household consumption by 0.8% in Russia. The insignificant wealth effect estimated for Ukraine accords with its relatively inefficient institutions, whereas, in Russia, the wealth effect seems to be indifferent to inefficient institutions due to the offsetting effect of the relatively large stock market.

JEL classification codes: C32, D12, E21, E44
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1. Introduction

The influential role of household consumption in recurring fluctuations of business cycles and determining standards of living has encouraged researchers to seek channels other than income that could affect consumer spending. A number of empirical papers have documented a positive relationship between values of houses and stocks, and household consumption. This finding has led to the introduction of new terminology: a housing wealth effect and a stock market wealth effect. A housing wealth effect implies consumption growth induced by an increase in the value of houses, while a stock market wealth effect implies consumption growth caused by an increase in the value of stocks.[1]

Along with the worldwide evidence on a causal relationship between stock market wealth and consumption, there is some evidence among developed economies that countries with better institutional settings tend to have larger wealth effects (Slacalek, 2006). In this light, we can suggest that the same regularity should also hold among emerging economies and especially in Eastern Europe (EE), since EE economies can generally be divided in two groups: one includes European Union (EU) members, which during the pre- and post-accession periods undertook substantial institutional reforms to meet EU standards, whereas the other group includes non-EU members, which did not have substantial incentives to reform. Therefore, we can expect that in contrast to the significant and large wealth effects found in EU EE countries (Vizek, 2011), the wealth effects in non-EU EE countries will be insignificant, or at least small. Extending the current literature to examine the association between wealth and household consumption in two non-EU EE countries, Russia and Ukraine, will allow us to conclude to what degree institutional environments matter for the significance and magnitudes of stock market wealth effects in EE emerging economies.

After the dissolution of the Soviet Union, the economies of its republics experienced industrial collapse, which led to significant declines in their GDPs and, as a consequence, forced households to decrease their consumption. The republics managed to recover only in the 2000s and, by the end of the decade, their economies outperformed their 1991 levels. Income growth accompanying economic recovery encouraged households to increase their spending.[2] The income growth was a key factor in the consumption growth, but not the only one. The other factor which could contribute to an increase in household consumption is stock market wealth growth. Although the financial crisis severely hit the stock markets in Russia and Ukraine, before the onset of the crisis their stock markets had demonstrated significant growth rates.[3] However, despite impressive growth rates of stock market wealth, the effectiveness of institutions in Russia and Ukraine remained low comparatively to other Eastern European countries (Figure 1).

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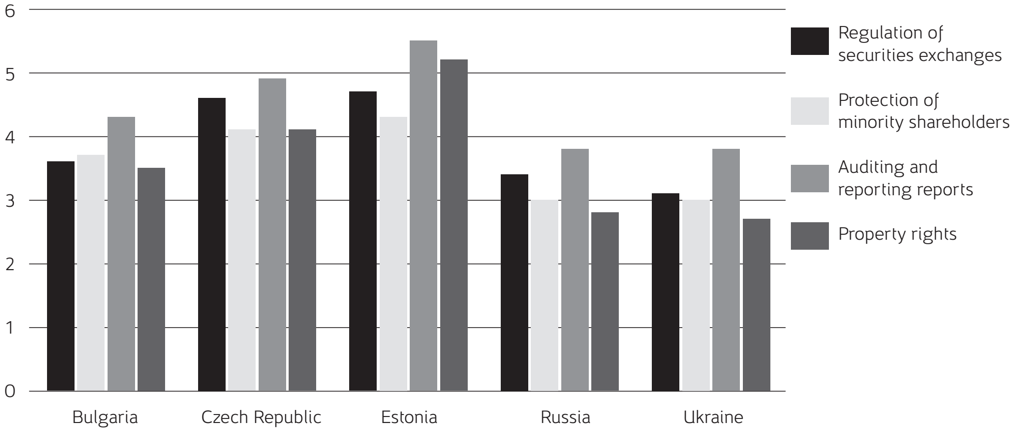
[1] Due to data limitations, this paper does not consider housing wealth effects.

[2] In Russia, household consumption in 2011 increased by 79% in real terms in comparison with the consumption level of 1995. Ukrainian households in 2011 consumed 77% more in real terms than they did in 1999. In the last decade, the share of household consumption in GDP was around 50% in Russia and 59% in Ukraine.

[3] The capitalisation ratio (market capitalisation as a proportion of GDP) in Russia grew 3.6 times within a decade, rising from 31.7% in 1997 to 115.6% in 2007. During the same period, the capitalisation ratio in Ukraine increased 10.7 times, reaching the value of 78.3% in 2007.

Figure 1. Indicators of Institutional Development, 2012



Note: 1=extremely weak and 7=extremely strong
Source: The Global Competitiveness Report 2012–2013.

The estimation results produced by the vector error correction models (VECM) show that in the long run, 10% growth in stock market wealth increases consumption by 0.8% in Russia, and by 0.1% in Ukraine; however, for Ukraine, the estimated wealth effect is statistically insignificant. While the low wealth effect for Ukraine could reflect the low institutional development, the large and significant wealth effect found for Russia undermines the initial hypothesis that inefficient institutions lead to insignificant and/or small wealth effects. Further analysis allows me to suggest that the large wealth effect found in Russia is due to the large size of its stock market.The article is organised as follows. The second section reviews the wealth effect literature. In the third section, the theoretical mechanism of a wealth effect is analysed, the wealth effect model is presented, and the relevant econometric methodology is introduced. In the following section, the employed data are described and their time series properties are examined. The fifth section analyses the estimation results. Finally, the last section presents conclusions.

2. Literature Review

The theoretical foundations of the wealth effect research rest on the life cycle theory of Modigliani and Brumberg (1954) and Ando and Modigliani (1963), and the permanent income theory of Friedman (1957), according to which households make consumption decisions based on expected lifetime resources. The econometric methods used to analyse consumption dynamics depend on the employed data structures. In the case of panel data, the usual choice is either the difference-GMM estimator proposed by Arellano and Bond (1991), or the panel co-integration techniques suggested by Pedroni (2000), Pesaran et al. (1999), and Mark and Sul (2003). In the analysis of time series data, the Engle and Granger (1987) two-step procedure, the VECM (Johansen, 1991, 1995), and the Dynamic OLS (DOLS) (Saikkonen, 1991; Stock and Watson, 1993) are usually employed. Studies also differ in terms of wealth definitions. The studies on wealth effects in developed economies employ direct measures of wealth, whereas the studies on wealth effects in emerging economies, due to reasons of data availability, use proxies for wealth. For example, stock market wealth is usually proxied by stock market indices.

To the best of my knowledge, only two papers have examined financial wealth effects in Central and Eastern European emerging economies. In the first study, applying the co-integration methodology to the quarterly data of four European emerging economies for the period from 1996 to 2010, Vizek (2011) finds statistically significant stock market wealth effects in all countries under consideration except Estonia. In particular, the author estimates that the elasticities of stock market prices are 0.12 in Bulgaria, 0.06 in Croatia, and 0.09 in Czech Republic. Furthermore, the net wage effect on consumption is significant only in Croatia and Czech Republic; the elasticities are 1.52 and 0.137 respectively.

The other study was conducted by Ciarlone (2011), who uses the quarterly unbalanced panel data for 17 emerging economies of Asia and Europe for the period 1995–2009. Using the pooled mean group estimator, the author concludes that when stock market prices rise by 10%, consumption increases by 0.2%. Furthermore, he finds that a 10% increase in disposable income results in a 7.6% increase in consumption. Additionally, Ciarlone (2011) makes separate estimations for a panel of the Asian emerging economies and a panel of the European emerging economies. While the disposable income elasticities for both Asian and Eastern European countries were almost the same, 0.72 and 0.71 respectively, the stock market price elasticity in Asian countries was two times larger than the elasticity in Central and Eastern European countries, 0.04 and 0.02 respectively. Finally, the author concludes that the adjustment to the long-run equilibrium goes faster in Asian economies than in Central and Eastern European economies.

There are also two papers which have studied the wealth–income association in non-European emerging economies. The first attempt to measure stock market wealth effects in emerging economies was made by Funke (2002), who examines the stock market price and consumption relationship in the annual panel data of 16 emerging economies over the period 1985–2000. Applying the fixed effects method, the author finds that a 10% increase in annual stock market returns leads to a 0.1–0.3% increase in private consumption.

Peltonen et al. (2009) apply the dynamic GMM method to the quarterly unbalanced panel data of 14 emerging economies for the period 1990–2008. The authors find that in the long run, a 10% increase in stock market wealth leads to a 1.28–1.48% increase in consumption, while a 10% growth in wages causes a 1.10–3.36% increase in consumption. Furthermore, for countries with long time series, Peltonen et al. (2009) run separate regressions. They conclude that financial wealth affects consumption significantly in all six countries. The reported long-run financial wealth elasticities are 0.05 for China and Taiwan, 0.09 for Hong Kong and Thailand, 0.11 for Korea, and 0.14 for Singapore. Furthermore, Peltonen et al. (2009) estimate that the wage elasticities are significant in all countries except Hong Kong, and their magnitudes vary from 0.241 for China to 0.623 for Thailand.

The wealth-consumption literature for developed countries is richer than that for emerging economies. For example, Case et al. (2001) apply the fixed effects method to the annual panel data of 14 OECD countries over the period 1975–1996, and find that the financial wealth and income elasticities are 0.02 and 0.66 respectively. However, when the authors control for either country-specific time trends or year-fixed effects, the estimated coefficients are statistically insignificant. In another study, using quarterly panel data for 16 OECD countries over the period 1985–2000 and the PMG estimator, Ludwig and Slok (2002) estimate that the elasticities of financial wealth and income are statistically significant and their magnitudes equal 0.08 and 0.70 respectively.Additionally, there is a series of papers which use time series data for wealth effect analysis. For example, Boone et al. (2001) find that financial and housing wealth have a statistically significant impact on consumption. In particular, the authors estimate that in the long run, the propensities to consume based on financial wealth range from 0.04 for the United States and the United Kingdom to 0.08–0.12 for France, Italy, Canada, and Japan. The estimates are obtained by applying VECM methodology to the quarterly data covering the period from 1970 to 1999. For Sweden, Chen (2006) applies the VECM framework to the quarterly data from 1980 to 2004, and concludes that the elasticities of disposable income and financial wealth are 0.427 and 0.06 respectively. Castro (2007) employs the DOLS method to estimate the wealth effect in Portugal; the dataset is quarterly and covers the period from 1980 to 2005. The estimated marginal propensity to consume based on financial wealth is 0.02, and the marginal propensity to consume based on disposable income is 0.61.

In conclusion of the literature review, it is worthwhile to note that the significance and magnitudes of estimated wealth effects for the same group of countries or even single countries can vary from one study to another. For example, the sizes of wealth elasticities found by Funke (2002) and by Ciarlone (2011) for emerging economies differ from those estimated by Peltonen et al. (2009). Such variations can occur due to differences in sample periods, variable definitions, estimators, and model specifications (Mehra, 2001). The differences in the magnitudes of stock market wealth effects among countries and country groups, in their turn, reflect the differences in sizes of stock markets, taxation, consumer behavior, wealth composition, market depth, volatility, stock market participation rates, and duration of participation (Funke, 2002; Hesse, 2008). Finally, in all studies, changes in stock market valuations have a modest effect on consumption compared to the effect of changes in disposable income. Poterba (2000) links such a small effect to a high concentration of wealth among a small number of households, bequest motives, and a willingness to accumulate precautionary savings.

3. Model and Methodology

The specification of the model is derived from the life cycle model proposed by Ando and Modigliani (1963). In this model, the consumer is supposed to maximise his or her utility subject to budget constraints, according to which, wealth at the end of the period equals gross savings plus interest income:

$$\max E_t \sum_{t=0}^{T} \beta^t u(c_t)$$

$$s.t.\ w_{t+1} = (1 + r)(w_t + y_t - c_t)$$

where w_t is the consumer's wealth at the beginning of the period, c_t is household consumption, y_t is current labor income, r is real interest rate, and β is a time preference factor.After applying several assumptions, such as a quadratic utility function, the equality of interest rate and time preference, labor income being an AR(1) process, and some algebraic manipulations, we derive the following specification, equation (1), which implies that current household consumption depends on its labor income and wealth:

$$c_t = \alpha y_t + \beta w_t \tag{1}$$

where α and β are the marginal propensities to consume based on income and wealth respectively.

In general, three transmission channels of stock market wealth can be distinguished: a realised wealth effect, an unrealised wealth effect, and a liquidity constraint effect (Ludwig and Slok, 2002). The realised wealth effect assumes that when stock holdings increase in value, households spend their gains, and, as a result, consumption rises. The unrealised wealth effect supposes that an increase in consumption is caused, not due to direct realisation of gains, but due to positive expectations regarding future wealth. This usually happens when households keep their stocks in locked-in accounts, such as pension accounts. Finally, the liquidity constraint effect implies that an increase in consumption is financed through increased borrowing supported by an increase in the value of households' portfolios.

Additionally, Poterba (2000) suggest that fluctuations in stock prices may also affect households that do not possess any stock, through the "confidence channel." Rising stock prices increase households' confidence about future economic conditions, and thus encourage them to increase spending. Declining stock prices, on the contrary, increase uncertainty about the future among consumers, and hence discourage consumption. The magnitudes of stock market wealth effects can vary among countries and country groups due to uncertainty regarding stock market returns.[4] Starr-McCluer (2002) and Rauning and Scharler (2011) find a negative relationship between stock market wealth effects and return uncertainty. The volatility in returns which create uncertainty among households can be caused by two types of risks. One type is risks that are specific to individual companies, and includes risks related to demand for products, litigation, resignation of top managers, the actions of regularity authority, nationalisation of assets, etc. The second type is economy-wide risks that affect every company in the economy, and, among others, includes liquidity risk, interest rate risk, exchange risk, tax risk, and credit risk. In fact, some of the mentioned firm-specific and economy-wide risks to a considerable degree, and the others to a certain extent, are influenced by institutional settings.

In this context, the central aim of institutions is to provide governance that will minimise risks, and therefore increase wealth effects. This aim is achieved through policies which require full disclosure of information on sold securities, prevent manipulation of price levels of securities, prohibit deceptive trading, facilitate liquidity, and limit "excessive" use of credit (Friend, 1975). Furthermore, recent empirical research supports the idea that high-quality institutions tend to reduce stock market volatility. Jayasuriya (2005) analyses the impact of stock market liberalization on stock market volatility in 18 emerging economies. The author finds that, in the post-liberalisation period, low volatility is observed in countries which have more developed institutions. Soo-Wah et al. (2012) examine the link between governance and equity market risk, and conclude that in emerging markets, there is a strong negative relationship between governance quality and equity market risk, whereas in developed countries, there is a weak or zero negative relationship.

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[4] Funke (2002) and Hesse (2008) indicate sizes of stock markets, market depth, taxation, stock market participation rates, and duration of participation as reasons for differences in the magnitudes of wealth effects. The factors mentioned in these studies depend on the decisions of households to engage in stock market activity, and the willingness of companies to be listed on the stock exchanges, which are influenced by the risks they will face while participating in the stock markets.

Equation (1) states that, in general, aggregate long-term consumption is determined by income and wealth. However, periodically, actual consumption may not be equal to long-term consumption. The reasons for such deviation include adjustment costs, habit persistence, and liquidity constraints (Mehra, 2001). Given that it takes some time for variables to return to the long-run equilibrium, I will estimate equation (1) using the VECM, which, in contrast to the other co-integration techniques such as fully modified OLS (FMOLS) (Phillips and Hansen, 1990), canonical co-integration regression (CCR) (Park, 1992), and DOLS, enables us to account for the short-run dynamics. Furthermore, in contrast to the FMOLS, CCR, and DOLS, the VECM allows for more than one co-integrating relationship. The VECM representation of equation (1) is the following:

$$\Delta z_t = \Pi z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta z_{t-i} + \varepsilon_t$$

where $z_t = \begin{pmatrix} c_t \\ y_t \\ w_t \end{pmatrix}$ and ε_t is the vector of innovations.

The rank of the coefficient matrix Π is the dimension of the co-integrating vector. If its reduced rank is $r<k$, where k is the number of variables, then $\Pi=\alpha\beta'$ where α and β are $k\times r$ matrices with rank r and βz_i is stationary. Each column of β is the co-integrating vector, and the elements of α are the error correction terms, which show how fast the dependent variable adjusts to deviations from the long-run state.

Johansen (1988) proposes a method to estimate the Π matrix and the likelihood ratio test, based on the trace statistic, or, alternatively, on the maximum eigenvalue statistic, to determine the number of co-integrating vectors. The trace statistic tests the null hypothesis of r co-integrating vectors against the alternative of k co-integrating vectors, while the maximum eigenvalue statistic tests the null hypothesis of r co-integrating vectors against the alternative of r+1 co-integrating vectors. However, before performing the likelihood ratio test, we need to make an assumption regarding the deterministic components. Johansen (1995) considers five different deterministic component cases:

- Case 1. No deterministic components
- Case 2. No deterministic trend in data, but intercepts in co-integration equations
- Case 3. Linear deterministic trend in data, intercepts in co-integration, and error correction equations
- Case 4. Linear deterministic trend in data, intercepts and linear trends in co-integration equations, and intercepts in error correction equations
- Case 5. Quadratic deterministic trend in data, intercepts and linear trends in co-integration and error correction equations

The results of the likelihood ratio tests, as well as the output of the VECM, are sensitive to the specification of the deterministic components. Johansen (1992) suggests a straightforward procedure to jointly determine the number of co-integration relations and specification of the deterministic components. According to this procedure, first we estimate all five specifications, and present the results from the most restrictive specification (Case 1 and r=0) to the least restrictive specification (Case 5 and r=k–1). Then we move from the most restrictive case to the least restrictive one, comparing the trace and/or maximum eigenvalue statistics to the corresponding critical values, and accept the case for which we reject the null hypothesis for the first time.

For the purpose of a robustness check, the consumption–wealth association will be re-estimated using the DOLS method. The DOLS representation of equation (1) has the following form:

$$c_t = \alpha_0 + \alpha_1 y_t + \alpha_2 w_t + \sum_{i=-k}^{l} \beta_i \Delta y_{t-i} + \sum_{i=-k}^{l} \gamma_i \Delta w_{t-i} + \varepsilon_t$$

where k and l indicate the number of lags and leads of the first differences respectively.

4. Data Description

To study the relationship between consumption and wealth in Russia, I use monthly data covering the period from January 2008 to December 2012, while in the estimation of the wealth effect in Ukraine, I employ quarterly data over the period quarter 1, 2002 – quarter 4, 2012. Different frequencies and time periods are used, only for reasons of data availability, and this imposes some limitations on the comparability of the results. The data on consumption and disposable income of the Russian households were taken from the database of the Federal State Statistics Service,[5] and the data on the stock market index were taken from the database of the MICEX Stock Exchange.[6] The source of the Ukrainian data is the database of the National Bank of Ukraine.[7]

Household consumption corresponds to total expenditures of households on goods and services, and income refers to disposable income, which is total income less taxes, social security contributions, and other mandatory payments. Household consumption and disposable income variables are expressed in per capita terms. As in previous studies, financial wealth is proxied by stock market indices expressed in domestic currencies. For Russia I use the MICEX index and for Ukraine the PFTS index. Furthermore, all variables are seasonally adjusted, converted into real terms by CPI index (all items), and measured in logs. Although in equation (1) the variables appear to be linear, I use logs of variables, because the logs of consumption, wealth, and income tend to be closer to linear values than do their levels (Ludvigson and Steindel, 1999).

Meanwhile, it should be noted that the choice of variables is subject to some limitations. First, household consumption includes both durable and non-durable consumption, whereas the theories consider only non-durable consumption because durable consumption is replacement of and addition to current stock (Mehra, 2001). However, Ludwig and Slok (2004) suggest that in stock market wealth effect studies, total consumption is preferable because financial crisis usually leads to reduction in durable, rather than non-durable, consumption. Second, I use disposable income, while conventional theories employ labor income. This substitution is to maintain the correspondence between consumption and income variables, because total consumption, rather than non-durable consumption, is used (Chen, 2006). Finally, the use of indices as proxies for stock market wealth makes it impossible to express wealth elasticities as marginal propensities to consume.

[5] http://www.gks.ru/wps/wcm/connect/rosstat_main/rosstat/en/main/
[6] http://rts.micex.ru/en/
[7] http://www.bank.gov.ua/control/en/index

Given that the study period includes the 2008 global financial crisis, I create dummy variables for this episode. The dummy variables are defined in the following way: they take the value of 1 from the period when consumption declines (the tenth month of 2008 for Russia, and the third quarter of 2008 for Ukraine) to the period when consumption returns to pre-crisis levels (the twelfth month of 2010 for Russia, and the fourth quarter of 2010 for Ukraine) and 0 otherwise.

Finally, in order to be able to apply the co-integration methodology, the variables must be integrated into order 1 (I(1)) processes. Therefore, augmented Dickey-Fuller (Dickey and Fuller, 1979) and Phillips-Perron (Phillips and Peron, 1988) tests are used to assess the time series properties of the variables. Both tests test the null hypothesis of the non-stationarity of series against the alternative hypothesis of stationarity. The ADF test is a parametric test whose results are very sensitive to the lag order. There are two rules for lag-order selection: an information-criteria based rule, and a general-to-specific sequential t test rule. The first rule assumes setting the lag order to the value which minimises information criteria (AIC and SIC). According to the second rule, we first set an upper bound for the lag length, and estimate the ADF test for this lag order. If the coefficient of the last lagged difference is statistically significant at the 5% level, then we accept the result of the test; otherwise we reduce the lag length by one, and repeat the procedure. If no lags are significant, we set the lag order to zero. In this study, the second rule will be used, since the simulations performed by Ng and Perron (1995) show that the sequential t test rule has comparable power and smaller size distortions. In contrast to the ADF test, the PP test uses a non-parametric method of controlling for serial correlation. It corrects serial correlation by directly modifying the test statistics.

The results of the tests, which are displayed in Table 1 and 2, show that in the case of Russia, the log differences of disposable income and financial wealth are stationary, implying that the variables are I(1), while consumption is trend-stationary. In the Ukrainian case, the log differences of all variables are stationary in all specifications of the unit root tests.

Table 1. Augmented Dickey-Fuller Test

			None			Constant			Constant and trend		
			Level	1st dif.	2nd dif.	Level	1st dif.	2nd dif.	Level	1st dif.	2nd dif.
Russia	c	test statistic	1,13	−5,83[a]	−6,45[a]	−2,27	−5,89[a]	−6,36[a]	−5,37[a]	−6,07[a]	−6,32[a]
		lag	2	6	6	0	6	6	4	6	6
		t(const)	–	–	–	2,27[b]	1,98	0,05	5,37[a]	−0,53	−0,43
		t(trend)	–	–	–	–	–	–	4,70[a]	1,58	0,49
	y	test statistic	0,99	−4,46[a]	−10,09[a]	−0,56	−4,55[a]	−10,00[a]	−2,55	−4,52[a]	−9,89[a]
		lag	2	1	1	2	1	1	3	1	1
		t(const)	–	–	–	0,57	1,00	−0,50	2,55[b]	−0,03	−0,27
		t(trend)	–	–	–	–	–	–	2,37[b]	0,60	0,05
	w	test statistic	−0,43	−4,61[a]	−10,59[a]	−2,73	−4,59[a]	−10,50[a]	−2,72	−4,61[a]	−6,27[a]
		lag	1	0	0	1	0	0	1	0	7
		t(const)	–	–	–	2,72[a]	−0,34	0,11	2,65[b]	−0,75	2,60[b]
		t(trend)	–	–	–	–	–	–	0,70	0,67	−2,23[b]
Ukraine	c	test statistic	5,41	−1,32	−9,70[a]	−1,33	−4,82[a]	−9,57[a]	−1,71	−4,89[a]	−9,45[a]
		lag	0	2	1	0	0	1	1	0	1
		t(const)	–	–	–	1,58	3,13[a]	−0,10	1,79	2,47[b]	−0,23
		t(trend)	–	–	–	–	–	–	1,36	−0,91	0,21
	y	test statistic	4,94	−4,07[a]	−7,87[a]	−1,60	−5,95[a]	−7,77[a]	−1,26	−6,20[a]	−7,64[a]
		lag	0	0	1	0	0	1	0	0	1
		t(const)	–	–	–	1,84	3,70[a]	0,03	1,36	3,22[a]	0,22
		t(trend)	–	–	–	–	–	–	0,87	−1,46	−0,23
	w	test statistic	−0,12	−3,46[a]	−7,13[a]	−2,12	−3,41[b]	−7,04[a]	−1,79	−3,58[b]	−6,95[a]
		lag	1	0	0	1	0	0	1	0	0
		t(const)	–	–	–	2,13[b]	0,16	−0,15	2,02	1,03	−0,02
		t(trend)	–	–	–	–	–	–	0,16	−1,07	−0,06

Test critical values

		None	Constant	Constant and trend			None	Constant	Constant and trend
Russia	1%	−2,61	−3,55	−4,12	Ukraine	1%	−2,63	−3,61	−4,21
	5%	−1,95	−2,91	−3,49		5%	−1,95	−2,94	−3,53

t statistics critical values

Russia	1%	2.66	Ukraine	1%	2.70
	5%	2.00		5%	2.02

Note: H0: a variable has a unit root; a: statistical significance at the 1% level; b: statistical significance at the 5% level.

Table 2. Phillips-Perron Test

			None			Constant			Constant and trend		
			Level	1st dif.	2nd dif.	Level	1st dif.	2nd dif.	Level	1st dif.	2nd dif.
Russia	c	test statistic	1,61	−11,47[a]	−22,32[a]	−2,02	−13,14[a]	−22,02[a]	−5,37[a]	−13,09[a]	−21,85[a]
		band	8	6	6	1	7	6	0	7	6
		t(const)	–	–	–	2,27[b]	0,88	0,28	5,37[a]	0,16	−0,15
		t(trend)	–	–	–	–	–	–	4,70[a]	0,30	0,32
	y	test statistic	1,48	−10,01[a]	−15,13[a]	−0,53	−10,19[a]	−14,94[a]	−1,99	−10,02[a]	−14,75[a]
		band	0	5	0	2	4	0	3	4	0
		t(const)	–	–	–	0,86	1,92	−0,38	2,14[b]	0,51	−0,60
		t(trend)	–	–	–	–	–	–	2,00	0,45	0,48
	w	test statistic	−0,57	−4,69[a]	−10,74[a]	−2,49	−4,67[a]	−10,64[a]	−2,42	−4,70[a]	−10,54[a]
		band	4	4	1	4	4	1	4	4	1
		t(const)	–	–	–	2.06[b]	−0,34	0,11	1,93	−0,75	0,07
		t(trend)	–	–	–	–	–	–	1,10	0,67	−0,02
Ukraine	c	test statistic	3,94	−3,42[a]	−11,06[a]	−1,18	−4,98[a]	−10,90[a]	−1,62	−5,03[a]	−10,74[a]
		band	4	5	3	4	4	3	4	4	3
		t(const)	–	–	–	1,58	3,13a	−0,20	1,41	2,47[b]	−0,22
		t(trend)	–	–	–	–	–	–	0,97	−0,91	0,14
	y	test statistic	4,42	−4,27[a]	−18,05[a]	−1,58	−5,97[a]	−17,82[a]	−1,32	−6,21[a]	−16,50[a]
		band	3	4	7	2	2	7	2	1	6
		t(const)	–	–	–	1,84	3,70[a]	−0,30	1,36	3,22[a]	−0,36
		t(trend)	–	–	–	–	–	–	0,87	−1,46	0,25
	w	test statistic	0,21	−3,51[a]	−7,29[a]	−2,06	−3,46[a]	−7,19[a]	−1,33	−3,67[b]	−7,09[b]
		band	3	2	3	3	2	3	3	1	3
		t(const)	–	–	–	2,05[b]	0,16	−0,15	1,39	1,03	−0,02
		t(trend)	–	–	–	–	–	–	−1,38	−1,07	−0,06

Test critical values

		None	Constant	Constant and trend			None	Constant	Constant and trend
Russia	1%	−2,61	−3,55	−4,12	Ukraine	1%	−2,63	−3,61	−4,21
	5%	−1,95	−2,91	−3,49		5%	−1,95	−2,94	−3,53

t statistics critical values

Russia	1%	2.66	Ukraine	1%	2.70
	5%	2.00		5%	2.02

Note: H0: a variable has a unit root; a: statistical significance at the 1% level; b: statistical significance at the 5% level.

In series with structural breaks, the traditional unit root tests can erroneously fail to reject the null hypothesis of non-stationarity (Perron, 1989). Since our sample periods include the global financial crisis of 2008, there is a possibility that the results of the ADF and PP tests are misleading. However, there is no need to conduct tests that allow for structural breaks, because the findings of the ADF and PP tests, that the variables are difference-stationary or trend-stationary, are in conformity with the results of the unit root tests of other studies (e.g. Vizek, 2011).

5. Empirical Results

As a first step before running the Johansen co-integration test (Johansen, 1991, 1995), an appropriate lag length has to be selected, because the results of the co-integration test are very sensitive to lag order. An appropriate number of lags for the Johansen procedure and, at the same time, for the VECM, is equal to the lag length of a well specified vector autoregressive (VAR) model (Sims, 1980) minus one lag. A VAR model is properly specified if its residuals are stationary and normal. The standard lag-order selection procedure identifies an appropriate number of lags as that at which information criteria are minimum. The choices of the lag-order selection procedure for the Russian and the Ukrainian models are displayed in Table 3.

Table 3. VAR Lag-Order Selection Criteria

Lag	Russia			Ukraine		
	AIC	SC	HQ	AIC	SC	HQ
0	−8.85	−8.62	−8.76	−3.16	−2.90	−3.07
1	−12.16	−11.59	−11.94	−9.72	−9.05	−9.49
2	−12.85	−11.94*	−12.50*	−10.16	−9.09*	−9.79
3	−12.93	−11.68	−12.45	−10.01	−8.54	−9.51
4	−12.84	−11.25	−12.24	−10.11	−8.24	−9.47
5	−12.91	−10.98	−12.17	−10.12	−7.85	−9.33
6	−13.01	−10.73	−12.14	−10.34	−7.67	−9.42
7	−13.03	−10.42	−12.03	−10.59	−7.52	−9.53
8	−13.25*	−10.30	−12.12	−11.11	−7.64	−9.91
9	−13.24	−9.94	−11.98	−12.88*	−9.01	−11.54*

Note: The optimal lag orders are indicated by an asterisk; AIC: Akaike information criterion, SC: Schwarz information criterion, HQ: Hannan-Quinn information criterion.

In the case of Russia, the AIC favors eight lags whereas the SC and HQ favor two lags. However, when the VAR model is estimated with either eight or two lags, the serial correlation LM test detects autocorrelation. When other lag orders are tried, it is found that only at the lag orders of four and six are the residuals of the VAR model both stationary and multivariate normal (Tables 4 and 5). Given that the AIC favors a model with six lags, subsequent analysis will be based on a six-lag VAR model.

Table 4. VAR Residual Serial Correlation LM Tests

Lag	Russia				Ukraine							
	VAR(4)		VAR(6)		AR(3)		VAR(4)		VAR(5)		VAR(6)	
	LM-stat	Prob	LM-stat	Prob	LM-stat	Prob	LM-stat	Prob	LM-stat	Prob	LM-stat	Prob
1	14.13	0.12	9.88	0.36	15.43	0.08	11.55	0.24	6.64	0.67	10.33	0.32
2	13.23	0.15	16.19	0.06	11.36	0.25	15.22	0.09	11.20	0.26	11.16	0.26
3	7.67	0.57	13.71	0.13	8.29	0.51	16.41	0.06	10.37	0.32	5.76	0.76
4	10.59	0.31	9.11	0.43	9.55	0.39	5.69	0.77	9.29	0.41	8.98	0.44
5	11.79	0.23	5.14	0.82	6.71	0.67	6.84	0.65	6.27	0.71	2.87	0.97
6	5.92	0.75	7.52	0.58	14.13	0.12	10.39	0.32	5.45	0.79	7.56	0.58
7	12.79	0.17	8.360	0.50	10.99	0.27	15.56	0.08	6.22	0.72	10.21	0.33
8	3.97	0.91	8.11	0.52	5.76	0.76	10.19	0.34	2.72	0.97	4.48	0.88
9	6.65	0.67	9.38	0.40	11.61	0.24	5.10	0.83	4.91	0.84	4.27	0.89
10	12.56	0.18	6.03	0.74	15.52	0.08	15.17	0.09	13.49	0.14	9.77	0.37
11	9.80	0.37	12.15	0.21	10.23	0.33	10.45	0.32	14.82	0.10	8.47	0.49
12	4.87	0.85	4.68	0.86	11.27	0.26	8.26	0.51	10.73	0.29	6.37	0.70

Note: H_0: no serial correlation at lag order h

Table 5. VAR Residual Normality Tests

	Russia				Ukraine							
	VAR(4)		VAR(6)		VAR(3)		VAR(4)		VAR(5)		VAR(6)	
	χ^2	Prob	χ^2	Prob	χ^2	Prob	χ^2	Prob	χ^2	Prob	χ^2	Prob
Jarque-Bera	5.24	0.51	4.90	0.56	4.22	0.65	2.99	0.81	3.69	0.72	6.06	0.42

Note: Cholesky (Lukepohl) is used for orthogonalisation, χ^2 – chi-square statistic, H_0: residuals are multivariate normal.

In the Ukrainian case, two information criteria give preference to nine lags, but the residuals of the VAR model with this lag order suffer from autocorrelation. Hence, the lag choice of the SC is tested. Although with this lag order the serial correlation and normality tests do not detect any problems in the residuals of the VAR model, the VECM produces implausible results. When the VAR model is estimated with other lag orders, it is found that with lag orders of three, four, five, and six, the model residuals do not have autocorrelation or non-normality problems (Tables 4 and 5). Among all lag-order specifications, the AIC favors the VAR model with six lags. However, for the VAR model with six lags, the estimated co-integrating equations also lack sensible interpretability; therefore the VAR model with five lags is selected as the second most favored model by the AIC.

After the proper lag orders for the Russian and Ukrainian VAR models are determined, the Johansen co-integration test is applied to check for the presence of co-integration relationships among the variables. The results of the test for both countries are given in Table 6. The likelihood ratio test results for Cases 1 and 5 are not reported, as they are highly unlikely cases. Case 1 is unlikely because the exclusion of the deterministic components does not allow us to account for the units of measurement. Case 5 is not considered because, when the logs of the variables are used, the inclusion of quadratic trends implies an always increasing or always decreasing rate of change (Harris and Sollis, 2003). Applying the Johansen (1992) procedure to the test results, one concludes that for Russia, the most appropriate model specification is Case 2 with two co-integrating vectors, while for Ukraine, Case 2 with one co-integrating vector is selected.[8]

Table 6. Johansen Co-Integration Test

		Case 2				Case 3				Case 4			
	No. of CE(s)	Tr	0.05	Max	0.05	Tr	0.05	Max	0.05	Tr	0.05	Max	0.05
Russia	None	79.86	35.19	47.34	22.30	73.80	29.80	45.98	21.13	86.34	42.92	47.39	5.82
	At most 1	32.52	20.26	26.79	15.89	27.82	15.49	24.75	14.26	38.96	25.87	25.75	9.39
	At most 2	5.73[b]	9.16	5.73[b]	9.16	3.07	3.84	3.07	3.84	13.21	12.52	13.22	2.52
Ukraine	None	43.56	35.19	25.12	22.30	29.66	29.80	20.39	21.13	41.01	42.92	21.02	5.82
	At most 1	18.45[b]	20.26	14.30	15.89	9.27	15.49	7.46	14.26	20.00	25.87	15.54	9.39
	At most 2	4.15	9.16	4.15[b]	9.16	1.81	3.84	1.81	3.84	4.46	12.52	4.46	2.52

Note: b: statistical significance at the 5% level, Tr –trace statistic, Lag orders: 5 (Russia) and 4 (Ukraine), Max – maximum eigenvalue statistic, 0.05 – the 5% critical value.

Finally, Table 7 reports the output of the VECMs for Russia and Ukraine. I report only the long-run equations, since the estimates of the short-run equations, in contrast to the long-run equations, can give us false inferences if consumption includes durable goods (Mehra, 2001). Since there is more than one co-integrating vector, restrictions on the co-integrating relations and the adjustment coefficients must be imposed to achieve identification. In the first co-integrating vector, which represents the consumption–income–wealth relationship, normalisation is achieved by setting the consumption coefficient to 1. As the second restriction, the second adjustment coefficient is constrained to 0. This constraint implies that deviations from the long-run income level, modeled in the second co-integrating equation, do not affect consumption in the long run. The theoretical foundation for this constraint can be found in the permanent income hypothesis (Friedman, 1957), which states that short-term fluctuations in income have no influence on consumption. In the second co-integrating vector, income is normalised while consumption is constrained. The wealth variable is not constrained, since it is assumed that a change in financial wealth, proxied by a change in stock market index, predicts a change in productivity (Levine and Zervos, 1998), and therefore predicts a change in income. Furthermore, the first adjustment coefficient is constrained based on the assumption that deviations from the long-run consumption level have no effect on the long-run income level. Finally, to check the validity of the restrictions, I use the LR test, whose resulting statistics confirms the validity of the imposed restrictions.

[8] It was expected that for Russia, Case 4 would be appropriate, since its consumption series appeared to be a trend-stationary process. However, the Johansen procedure selected Case 2. This result is not surprising, since it can be the case that a consumption series, in fact, is a unit root process rather than a trend-stationary process, because in finite samples, a unit root series can be approximated by a trend-stationary series and vice versa due to the closeness of autocovariance structures (Harris & Sollis, 2003). As a result, the probability of a false rejection of the null hypothesis of a unit root process increases.

Table 7. Estimation Results of Vector Error Correction Models

Variable	Russia		Ukraine
Co-integration equations			
C	1.00	0.00	1.00
Y	−0.69[a] [−13.66]	1.00	−0.93[a] [−25.99]
W	−0.08[a] [−3.72]	−0.26[b] [−2.14]	−0.01 [−0.56]
Constant	−2.61[a] [−4.89]	−7.54[a] [−8.90]	−0.41[b] [−2.03]
Error correction terms			
Term 1	−0.99[a] [−4.03]	0.000000 [0.00]	−0.39[b] [−2.09]
Term 2	0.00 [0.00]	−0.13[a] [−4.87]	– –
Restrictions			
Co-integration restrictions	Yes		No
LR test for binding restrictions			
Chi-square	3.65		–
Probability	0.07		–
No. of observations after adjustment	54		39
Lag intervals	5		4
Rank	2		1
t statistics critical values			

Russia	1%	2.66	Ukraine	1%	2.70
	5%	2.00		5%	2.02
	10%	1.67		10%	1.68

Note: *t* statistics are in brackets, [a] statistical significance at the 1% level, [b] statistical significance at the 5% level, [c] statistical significance at the 10% level.

The co-integration equation shows a statistically significant positive relationship between financial wealth and consumption in Russia. The magnitude of the financial wealth coefficient implies that a 10% financial wealth increase results in a 0.8% increase in consumption. Furthermore, the coefficient for the income variable also appears statistically significant at the 5% level and suggests that, for every 10% increase in disposable income, consumption increases by 7%. The large value of the adjustment term implies a quick adjustment of consumption to deviation from the long-run level.

In the case of Ukraine, the coefficient of the wealth variable implies that a 10% increase in financial wealth induces a 0.1% growth of consumption; however, the effect is not statistically significant. Table 7 also shows that a 10% increase in disposable income is predicted to raise consumption by 9.3%. Furthermore, the adjustment coefficient is significant at the 5% level, and its size indicates that households adjust to deviation from the long-run level approximately within two and a half quarters.

It also should be noted that the critical values of the Johansen test do not take into account exogenous variables. Hence, the presence of crisis dummy variables in the Russian and Ukrainian models can make inference problematic. However, there is only a risk of underreporting the true number of co-integrating vectors, since simulations run by Johansen and Nielsen (1993) show that the inclusion of an exogenous dummy variable leads to lower critical values. This risk did not affect the Russian model, since the test finds the maximum possible number of co-integrating vectors. However, in the case of Ukraine, the test may underreport the true number of co-integrating relations; therefore, the VECM was re-estimated with two co-integrating vectors. However, the results did not change dramatically, leaving the inference of the VECM with one co-integrating vector unaltered.

To verify the robustness of the estimated parameters, the wealth–consumption relationship was estimated using the DOLS method. As with any estimation method which involves lags and/or leads, DOLS results are also sensitive to lag and lead orders. The appropriate lag and lead orders are selected using an AIC. By restricting the maximum number of leads and lags to six, in order to avoid the danger of over fitting, I find that the AIC favors five lags and six leads for the Russian model, and four lags and four leads for the Ukrainian model. In addition, the ADF test without exogenous terms shows that, with given lag and lead orders, the residuals of the DOLS models of Russia and Ukraine are both stationary. The other specifications of the ADF test are not employed, since the DOLS models already include the necessary deterministic terms. The DOLS estimates are displayed in Table 8.

Table 8. DOLS Estimates

	Russia	Ukraine
Y	0.70[a] [5.23]	0.90[a] [38.45]
W	0.13[c] [1.82]	0.01 [0.67]
Constant	2.26[b] [2.01]	0.64[a] [4.63]
Lags	5	4
Leads	6	4
No. of observations after adjustment	48	35
R^2	0.89	1.00

t statistics critical values

Russia	1%	2.66	Ukraine	1%	2.70
	5%	2.00		5%	2.02
	10%	1.67		10%	1.68

The stationarity test on residuals
ADF test without exogenous terms

t statistics	−8.63[a]	−4.69[b]
Lag	0	5
Critical values		
1%	−4.61	−4.73
5%	−3.92	−3.99

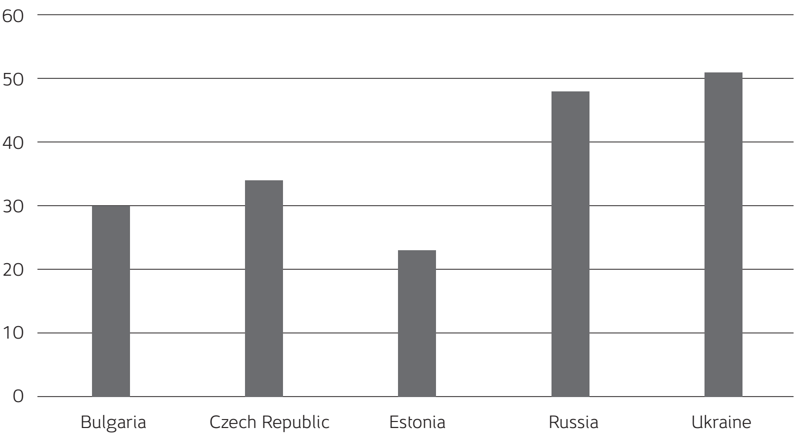
Note: Dependent variable: c, HAC standard errors and covariance, t statistic in parentheses, lead and lag orders based on AIC, critical values for the stationarity test are based on MacKinnon (1991), t statistics are in brackets.
[a] statistical significance at the 1% level, b statistical significance at the 5% level, c statistical significance at the 10% level.

The DOLS estimate of the income effect in Russia is still significant at the 5% level and its magnitude is equal to that of the VECM estimate. Furthermore, while the VECM coefficient of stock market wealth is statistically significant at the 5% level, the DOLS coefficient is significant at the 10% level and the size of the wealth coefficient is greater by 0.05 points. The reason that the DOLS method produces results different from those of the VECM lies in its technical characteristics. Thus the DOLS approach assumes one co-integrating relationship, and, in the case of Russia, this assumption may be problematic, because the likelihood ratio test detects two co-integrating vectors. On the contrary, in the case of Ukraine, for which the likelihood ratio test finds one co-integrating vector, the DOLS method produces results qualitatively and quantitatively similar to those of the VECM.

The insignificant stock market wealth effect found for Ukraine accords with our expectation that ineffective institutions, through higher uncertainty, lead to lower wealth effects (Figure 2). However, the finding that in Russia, the wealth effect is statistically significant, and its magnitude is close to those for Bulgaria and Czech Republic in Vizek (2011) despite comparatively higher uncertainty, runs counter to our anticipation that ineffective institutions lead to lower wealth effects.9 This finding implies that perhaps there exists another factor, specific to the Russian economy, which outweighs the negative impact of inefficient regulation. Hence, the most likely reason for the large wealth effect estimate must be sought among the factors that also depend on influences other than institutions.

Figure 3 allows us to conclude that the size of the Russian stock market, which surpasses the other regional stock markets, could be responsible for the large wealth effect, because markets with greater capitalisation tend to produce larger wealth effects (Funke, 2002). The large capitalisation of the Russian market is not surprising because the majority of listed companies represent the world's largest producers of oil, metals, and chemicals, and, hence, weak institutional development carries little cost to them.

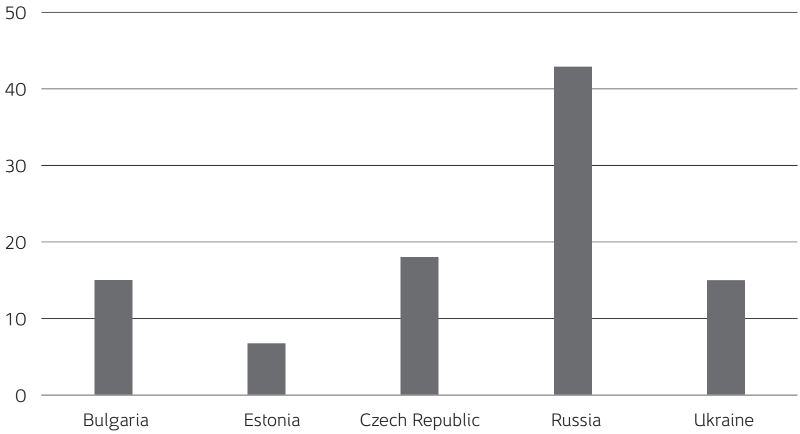
Figure 2. Volatility of Stock Price Index, 2010



Note: Volatility of stock price index is the 360-day standard deviation of the return on the national stock market index.
Source: The Little Data Book on Financial Development, 2013.

[9] The comparison should be taken with some skepticism due to different time periods and frequencies.

Figure 3. Stock Market Capitalisation, % of GDP, 2011



Source: The World Bank

7. Conclusions

Empirical results suggest that in Russia, changes in both stock market prices and disposable income have statistically significant effects on consumption, while in Ukraine, only changes in disposable income have a significant impact on consumption. According to the VECM results, a long-run increase in stock market prices by 10% increases household consumption by 0.8% in Russia and 0.1% in Ukraine, whereas the long-term elasticities of disposable income are 0.70 in Russia and 0.90 in Ukraine.

On the whole, these findings imply that, as in other countries, consumption behavior in Russia and Ukraine is mostly driven by disposable income, rather than stock market wealth. Therefore, transfers and taxes remain the main tools of policy makers in stimulating consumption. Meanwhile, since the wealth effect in Russia is significant, policy makers should closely watch developments in the stock markets, and prevent excessive fluctuations in stock prices through monetary policy and financial regulation, in order to limit the negative impact of declines in stock market wealth on aggregate demand. Furthermore, given that institutions are relatively weak, policy makers can strengthen the stock market wealth channel by tightly bridging the gap between the existing securities and related legislation, and the standards promoted by the Internal Organization of Securities Commissions (IOSCO).

The insignificant stock market wealth effect found in Ukraine should in no way imply that policy makers can neglect this channel; instead, they need to enhance this channel and then use it as a supplementary one. For this purpose, Ukrainian policy makers need to take significant steps toward the convergence of stock exchange and related legislation with IOSCO standards. However, full convergence of regulation with the highest standards will not alone yield the desired effect if not supported by a favorable macroeconomic environment, which will also contribute to the moderation of volatility.

As for further research, it would also be interesting to examine the possibility of non-linear stock market wealth effects in Russia and Ukraine. Such a research interest is driven by the hypothesis that, as positive and negative news influences stock market performance asymmetrically, positive and negative changes in stock market wealth also affect household consumption differently. For example, Apergis and Miller (2006) find that, in the U.S.A., positive changes have a greater impact on consumption than negative changes. Additionally, it would be useful to conduct research on wealth effects using micro-level data, in order to verify the validity of the “confidence channel” hypothesis in EE countries.

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