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CONTENTS

Foreword

Actions of "economic agents" are, in addition to being driven by "economic" incentives, affected by mood and emotions, as well as by other "non-economic" incentives. Prior to the "outbreak" of the last global financial and economic crisis, we experienced a relatively long period of growth and increase in prosperity, observed through increases in purchasing power and/or increased government spending. Extremely positive mood and emotions, mostly resting on two important traits of humans - salience and recency, led to a gross underestimation of "risk" and a gross overestimation of future "returns". In the aftermath of the global financial and economic crisis the tables have turned - mood and emotions have become extremely negative and sentiment is essentially determined by the perception of "risk". It is therefore not surprising that these days most of the talk, either academic, regulatory, professional, home or bar, revolves around risk (albeit in implicit terms). Financial industry does not discuss asset allocation any longer, as professionals prefer to talk about risk allocation or risk parity. Regulators are concerned with fraud risk and related financial reporting standards, with systematic risk borne by (too-big-to-fail) financial institutions and herding behavior of investors, especially bank depositors. Individuals worry about their future, about the riskiness of their savings placed either in bank deposits or entrusted to pension funds.

The common theme of the papers published in this issue of Economic and Business Review is risk (albeit sometimes in implicit terms). In the first paper Berk Skok, Cok, Košak, and Sambt show the negative effects of demographics dynamics on the future pension benefits based on the pay-as-you-go system in Slovenia. Authors then continue to demonstrate benefits of private retirement savings and the important role of asset allocation decision in terms of time horizon and equity allocation. In the second paper Foye, Mramor, and Pahor investigate whether stock markets in Eastern European EU member states are weak form efficient. They show that the markets are not weak form efficient and provide possible explanations for such result. These findings have important implications for financing and investment policies of firms. In the third paper of this issue Asanovic explores the determinants of the so-called early warning models for systemic banking crisis in case of Montenegro. The author performs the analysis using Bayesian model averaging, which minimizes the subjective judgment related to the choice of early warning indicators. Finally, in the last paper of the issue Koren and Valentincic analyze to what extent do U.K. public companies engage in earnings management in order to meet dividend payout thresholds. Their findings suggest that dividend payout thresholds are significant determinants of earnings management behavior, which confirms the important role of dividends as perceived signalling mechanism.

Igor Loncarski,
associate editor

THE ROLE OF ASSET ALLOCATION DECISIONS IN PLANNING FOR A PRIVATE PENSION: THE CASE OF SLOVENIA

ALEŠ BERK SKOK[1]
MITJA ČOK[2]
MARKO KOŠAK[3]
JOŽE SAMBT[4]

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ABSTRACT: *Current demographic dynamics driven by low fertility and increasing longevity requires adjustments of the traditional frameworks of providing pensions. In this article we highlight three crucial issues policymakers should address by implementing those adjustments. First, fiscal limitations given the current and projected demographic dynamics will dramatically reduce PAYG pensions. Without sufficient savings during the active period, individuals will increasingly end up in poverty. Their savings will not be enough to support their desired consumption in old age. Second, we highlight the impact of the asset allocation decision and the general public's related lack of awareness on this issue. Therefore, we argue that financial illiteracy about both required savings and about decisions on appropriate asset class play a significant role in determining the well-being of masses in the not-so-distant future. Third, we argue that shift towards private pension away from the PAYG is expected to come with substantial benefits stemming from diversification among conceptually different sources of pension income.*

Key words: *PAYG, private pensions, financial literacy, old-age income, risk diversification, transition economics*
JEL Classification: J14, G11

1. INTRODUCTION

Population aging requires that the traditional pay-as-you-go (PAYG) systems are downscaled. Projections of age-related expenditures from the European Commission (DG ECFIN) and Economic Policy Committee (AWG) (2009) point toward a significant risk to the sustainability of PAYG systems as a consequence of increasing demographic shifts. Muenz (2007) argues that until 2050 demographic dynamics are pro-

[1] University of Ljubljana, Faculty of Economics, Slovenia, Ljubljana, e-mail: ales.berk@ef.uni-lj.si
[2] University of Ljubljana, Faculty of Economics, Slovenia, Ljubljana, e-mail: mitja.cok@ef.uni-lj.si
[3] University of Ljubljana, Faculty of Economics, Slovenia, Ljubljana, e-mail: marko.kosak@ef.uni-lj.si
[4] University of Ljubljana, Faculty of Economics, Slovenia, Ljubljana, e-mail: joze.sambt@ef.uni-lj.si

jected to result in a 10-year increase in the median age of the EU population, from 38 to 48 years old. Governments should build substantial funded pension systems as a supplement to the traditional PAYG (Du et al., 2011). More and more weight should be given to private pension systems, as under unfavourable demographic dynamics, they are far more efficient than PAYG systems (i.e., under realistic assumptions, they can deliver higher pension benefits with the same level of contributions or the same level of pension benefits with a lower level of contributions; Garrett and Rhine, 2005; Berk and Jasovic, 2007).

This long-term shift toward funded private pensions should be based on sound second- or third-pillar[5] frameworks, or both (Boersch Supan et al., 2008). Namely, trends in redesigning pension systems have during the past decade favoured the diversification of risks across all sources of old-age income as the coexistence of the three pillars positively effects benefits and consumption under various shocks, e.g. ageing population, inflationary shock, stock market crash etc. (World Bank Pension Conceptual Framework, 2008; Holzmann and Hinz, 2005; Lindbeck and Persson, 2003, Du et al., 2011). When a society is decreasing reliance on the PAYG and increasing reliance on private pension pillars, nature of co-movements between the drivers of pension benefits in both systems are of a great importance. Those co-movements can be measured with correlation coefficients between wages (predominant driver in the PAYG) and financial variables, i.e. stock and bond returns. Holzmann (2002) is the first published peer-reviewed research reporting very beneficial (i.e. low) national level correlation coefficients between wages and interest rates, and wages and capital return. Namely, he reports coefficients of correlation between wages and interest rates in the range between -0.197 and 0.238 and correlation coefficients between wages and capital return in the range between -0.077 and 0.202. Other authors in the area of diversification benefits report similar figures, in some cases even more beneficial (e.g. see Knell, 2010).

Despite the evident shift toward private pensions, one should not expect overnight changes. Augusztinovics (2002) argues that countries, even though they redesign their pension system and move to strengthen individual pension accounts, will still deliver their pensions predominantly from PAYG systems for quite some time. Recent experience of some countries in Central and Eastern Europe provide ample evidence of the budget constraint posed by high transition costs for cases of accelerated reform towards private pensions (Simonovits, 2011). Ferber and Simpson (2009) also argue that market meltdowns make shifts towards funded pillars less politically feasible. At the same time, private pensions should not be taken for granted, as only well-managed, efficient frameworks (e.g., competitive institutions and products, broad population coverage, sound governance mechanisms) can deliver the anticipated advantages (Pensions at a Glance, 2009; Bertranou et al., 2009).

[5] Second pillar includes mechanisms through which employers make contributions for their employees and third pillar the ones through which employees make their own contributions, regardless of the level of obligation.

The resulting pension landscape will not only provide a more sustainable and efficient environment for managing inter-temporal consumption but also support domestic underdeveloped financial markets. Davis (2008) shows that pension-fund growth in the European Union is likely to lead to beneficial financial development with a broader range of instruments and a lower cost of capital, thus leading to higher welfare. He further argues that pension-fund growth has a significant effect on Eurozone financial markets, by moving them partly toward the Anglo-American system, as well as promoting integration. Davis and Hu (2008) provide evidence that funding improves economic performance sufficiently enough to generate resources to meet the needs of an aging population and that the improvement is even greater in emerging market economies.

However, the previously mentioned characteristics of private pension systems are by themselves insufficient to provide for society's well-being if people do not have sufficient financial knowledge, i.e. are only modestly financially literate. Financial illiteracy is a very important issue, and it has been reported even for the most advanced countries (on the United States, see Lusardi and Mitchell, 2007; on the United Kingdom, see Gathergood and Disney, 2011; on Japan, see Sekita, 2011; on Germany, see Buchner-Koenen and Lusardi, 2011). Studies have found that many households are unfamiliar with even the most basic economic concepts in order to make savings and investment decisions. Financial illiteracy is lowest among women, young people, and individuals with lower incomes and lower education levels. With respect to pension savings, financial literacy increases individuals' likelihood of having a savings plan for retirement, which has a very strong impact on their wealth levels at retirement (Lusardi and Mitchell, 2007a).

We argue that very important aspect of financial literacy addresses knowledge about characteristics of various asset classes for their investments. Rooij et al. (2007) found that financially illiterate individuals are significantly less likely to invest in stocks. We show in this paper that this aspect has a very significant impact on the level of pension wealth, since choosing appropriate asset classes is extremely important. Strategic asset allocation determines approximately 90 percent of portfolio performance (see Brinson et al., 1986; Ibbotson and Kaplan, 2000; Andreu et al., 2010). Overall, it is crucial that financial literacy campaigns address both topics: individuals' need to start saving for their pension (e.g., in a pension savings account) and at the same time they also need to allocate savings into appropriate asset classes.

We focus on Slovenia, a country with a combination of a significantly aging population and an underdeveloped private pension system. Exclusive dataset on the distribution of individuals' income in Slovenia is used in this article to support our three main points, which are particularly important for people in countries like Slovenia who are entering a career or are halfway into their professional career. We contribute to the literature with the model, which shows the required monthly savings under each of three asset allocation choices (i.e., stocks, treasury bonds, and treasury bills). We calculate the required savings during the active work period of individuals' life under the assumption that they (together with the assumed long-term yield) can fill the gap between projected pensions from the PAYG system and the 70% net replacement rate suggest by the Organisation for Economic Co-operation and Development (OECD, 2009a). Different income levels (decile groups) are taken care of and insights into the potential outcome of a risk-aware individual allocating all of his or her pension savings into a risky diversified stock portfolio and prepares for poorly performing financial markets but actually achieving the long-term mean yield are offered. This case clearly favours investing in stocks over the long run. Finally, we address the issue of pension income diversification and show that benefits are greatest at the point, where private pension pillars only start to provide pension income. Our conclusions are relevant in general, i.e. for many developed countries across the globe, as not many current pension systems have sufficient solutions regarding an increasing old-age dependency ratio.

This article is structured as follows. In the second section, we briefly describe the existing Slovenian private pension system and present pension funds in the context of the whole financial market. We also report the performance of Slovenian pension vehicles since their introduction nearly a decade ago and compare that with the performance of pension funds from developed markets. In the third section, we describe benefits from the Slovenian PAYG system and related taxation. The fourth section offers demographic projections up to 2060 and future public pension expenditures, which without changes, are expected to cause huge deficits in the pension budget. As those imbalances are unsustainable and cannot be financed through subsidies from the central government budget, we impose fiscal caps at various percentages of gross domestic product (GDP) that can be allocated to finance pensions. Those in turn pose further caps on the future levels of expected public pensions. The fifth section provides overview of three basic asset classes available for the allocation of private pension savings. Using historical data, we calculate real long-term yield and further assume that those returns are a reasonable approximation of future long-term yields. We thus use historical returns as expected returns in our model, which we present in detail in the sixth section. In section seven we present the extent diversification benefits.

2. SLOVENIAN SYSTEM OF PRIVATE PENSIONS

The pension reform enacted in Slovenia in 2000 introduced private pensions within the second pillar, which comes in two forms. The first form are pensions, which are compulsory for employees in "health-risk" jobs. Employers must make special pension contributions for all such classified workers, and those contributions are transferred to employees' pension account at the special pension fund (managed by a government-sponsored institution). Second, for all other employees, participation in the defined contribution pillar is not compulsory but is promoted by a tax incentive. Namely, contributions to the second-pillar pension funds are subject to tax relief at the level of a payer. Either an employer or an employee can make a contribution, but the total amount of tax relief cannot surpass either the maximum of 5.844% of an employee's annual gross wage or a cap that is set annually[6] When an employer pays a second-pillar contribution for employees, it can deduct paid contributions from the company's corporate income tax base, while in the case that a second-pillar contribution is paid by an employee, it is deducted from her personal income tax base.

Table 1: *Size of the second pillar, average contribution to the second pillar and breakdown of total assets at the end of 2012*

	MPFs	PCs	ICs	Total
AUM (mln EUR)	839.0	655.0	302.6	1,796.6
Average annual contribution	450.72	466.92	381.47	422.53
Breakdown of total assets (%)*				
Deposits	22.2	22.7	n.a.	n.a.
Government bonds	28.1	37.4	n.a.	n.a.
Bonds: other	29.1	32.6	n.a.	n.a.
Stocks	1.1	5.3	n.a.	n.a.
Investment funds	19.0	0.0	n.a.	n.a.
Cash	0.5	2.0	n.a.	n.a.
Total assets	100.0	100.0	n.a.	n.a.

Note: MPFs = mutual pension funds, PCs = pension companies, ICs = insurance companies, and AUM = assets under management; * - PC breakdown of total assets at the end of 2011.
Sources: Ministry of Labour, Family and Social Affairs (http://www.mddsz.gov.si), Report on financial market trends (2013); Report on insurance market trends (2012).

There were 508 thousand participants in the second pillar by the end of 2012, which represents 60.7% of the total number of persons in employment[7]. Different second-pillar institutions manage the individual pension accounts: insurance companies (ICs), pension companies (PCs), and mutual pension funds (MPFs). At the end of 2012 total assets under management of the second-pillar institutions was only 1,797 mln EUR, as the average annual contribution is only about 400 EUR. Assets represented only 2.1% of the assets of the overall financial sector and only 5% of the GDP (Bank of Slovenia, 2013).

A notable characteristic of the Slovenian private pension system is inappropriate asset allocation. Rules about guarantees in the private pension system (Pravilnik o izračunu…, 2005) force pension managers to reach a certain percentage (at least 40%) of the cumulative yield of long-term bonds issued by the Treasury of the Republic of Slovenia on a single-member contribution. Because pension asset managers must provide additional capital in the case that their products don't deliver the guaranteed threshold yield, they do not take much risk. As a result, they tend to invest less than 5% in stocks, even though participants in the pension fund might have investment horizons extending as far ahead as 40 years. Fixed-income instruments together with

[6] The cap was 2,526.2 EUR in 2008, 2,604.5 EUR in 2009, 2,646.2 EUR in 2010, 2,683.3 EUR in 2011 and 2,755.71 EUR in 2012 (Tax Administration, 2012).

[7] 62% we obtain using the registry data on number of persons in employment. Using the definition of International Labour Organization (ILO) the share is even lower (54%).

deposits and cash represent at least 90% of total assets (see Table 1). Asset allocation in developed countries is dramatically different, as stocks represent roughly half of the total assets allocated.[8]

Of course, ultraconservative asset allocation can yield only meagre performance. When portfolio strategists set a conservative floor for the portfolio, the ceiling is not very high (Jensen and Sorensen, 2001). In the period 2003-2012 Slovenian pension funds recorded only 1.05% average real annual yield (mutual pension funds [MPFs]) and 0.87% (pension companies [PCs]).[9] Pension vehicles as a group beat pension guarantee, as in real terms guarantee only amounted to –0.60%. Figure 1 shows the dynamics of the real yield of MPFs, PCs, the private pension system guarantee, and the best and the worst performer (in the entire 2003-2012 period) of all products in the market for the period 2003–2012 in Slovenia.

If we compare same-period performance of Slovenian pension products with similar products in developed countries, we see that those countries did not have much better performance. However, there is a conceptual difference between Slovenian private pension products and those in the developed world. It is impossible to achieve long-run performance of 6–10% typical for countries with the developed private pension systems[10] with the strategic asset allocations of Slovenian pension products, all of which are characterized by investment policy unification regardless of the age of their members and all of which are ultraconservative.

Because Slovenia's private pension system cannot offer appropriate savings vehicles, it should change and pension products with less conservative exposure should be offered.[11] Under a new system, individuals should have their choice of asset allocation—individuals have different characteristics and needs, and not all of them need a guarantee. In the "Results" section, we point out the significant impact of the asset allocation decision on the outcomes of pension savings.

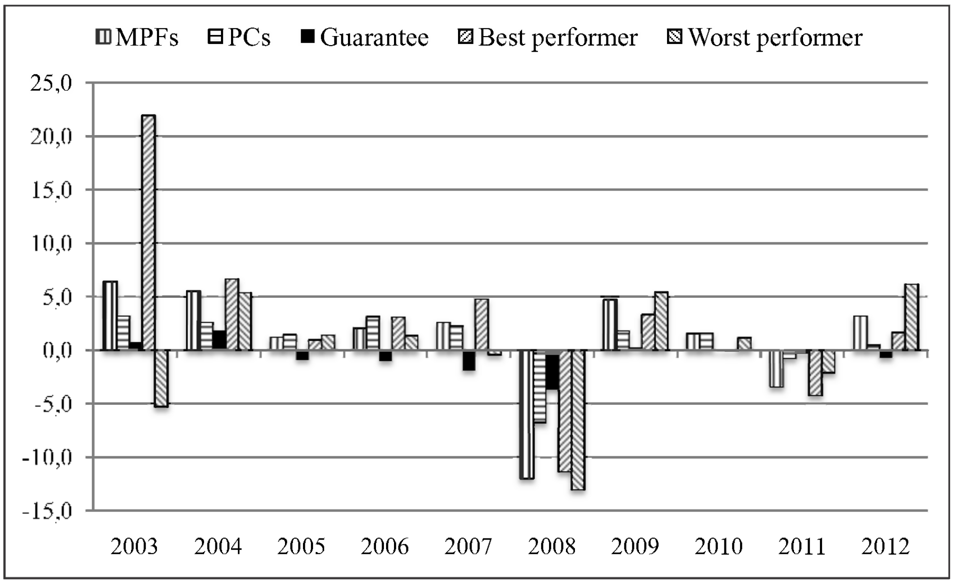
[8] For the end of 2009, a Towers Watson study reported the following stock allocations: Australia, 57%; Canada, 49%; Hong Kong, 62%; Japan, 36%; Netherlands, 28%; United Kingdom, 60%; and United States, 61% (2010 Global Pension Asset Study, 2010).

[9] Comparison of yields between MPFs and PCs must be taken with a grain of salt, as PCs are allowed not to mark to market all of their assets.

[10] Antolin (2008) reports performance between 6 and 8 percent in real terms since introduction of private pension systems, measured in geometrical terms.

[11] The legislation, which allows for asset-allocation investment policy design, became effective on Jan 1 2013 (Pension and Disability Insurance Act, 2012), but the second level rules and pension products are still being prepared.

Figure 1: *Dynamics of real annual yields of MPFs and PCs in the period 2003–2012 in Slovenia (in %)*



Source: Report on financial market trends (2013), Report on insurance market trends (2012), Monthly Bulletin (2005, 2009, 2013), authors' calculation (averages within MPFs, PC categories).

3. OVERVIEW OF CURRENT BENEFITS FROM THE SLOVENIAN PAYG SYSTEM

Under the pension law being is in force from 1 January 2013 the total accrual rate for an individual with full retirement conditions is 57.25%. By assuming his wage was growing in line with the average wage in Slovenia net replacement rate amounts to 57.25% as well. The pension base is calculated as the average from the individual's valorised best consecutive 18 years (19 in 2013, 20 in 2014 and finally 24 in 2018 and onwards).

Individual's gross wages by years are transformed to nominal wages with the ratio between average net and gross wage in that year. Those 'net' wages are multiplied with the vector of valorisation coefficients[12] to calculate the pension base. Finally, accrual rate is applied to the pension base to calculate the amount of first pension. Accrual rate for men amounts to 26% for the first 15 working years and further 1.25% for each additional working year. Thus, for a man with 40 working years total accrual rate is 57.25% (26% + 1.25% * 25 years). For women the pension system is more generous with 29% for the first 15 years and, again, 1.25% for each additional working year. Thus, for women with

[12] Calculation of those factors was based on the past growth of pensions relative to wages. After the pension law introduced with 1 January 2013 the set of factors will grow in line with the growth of average wage – thus, it will not depend on the growth of pensions any more. In the past factors were declining because the pension growth was lagging behind the wage growth.

40 working years the total accrual rate of 60.25% is applied. From 2013 to 2022 there is a transition period in which less than 40 working years is required for women and therefore in this period for women even higher accrual rate is applied for 40 working years[13]. However, in our calculations we focus only on male with 40 working years.

Slovenia, as other countries from Central and Eastern Europe, has undergone through several phases of pension reform; the last phase had passed the Parliament in 2012 and is effective from January 2013. Its most important element is a gradual increase of retirement age for both genders. The full retirement age (for old age pension) is thus increasing from 61 years (women) and 63 years (men) to 65 years (by 2016 for men and by 2020 for women). However, under both pension systems earlier retirement (up to several years) was/is possible with full benefits and without penalties if the person collects required number of working years earlier. Slovenia is also characterised by the fact that it has never implemented a compulsory second pillar[14] compared for example with Slovakia, Hungary and Poland. However, the compulsory second pillar has been recently effectively abandoned in Hungary, while in Slovakia it is not compulsory any more since February 2013. The Czech Republic which initially also did not introduce mandatory second pillar is now opening the option for employees to divert part of their contributions from the first to the second pillar (Berk et al., 2013).

3.1. Financing the PAYG pillar

Compulsory pension contributions for the PAYG pillar are set at the rate of 24.35% (employees pay 15.5%; employers, 8.85%) out of a gross wage without any ceiling.[15] The aggregate contributions total 3,348.9 million EUR, or 9.5% of GDP, in 2012. Because this is not sufficient to cover expenditures of the first pillar (which totalled 4,851.0 million EUR, or 13.7% of GDP, in 2012) in aiming to maintain the financial stability of the system, current legislation has stipulated that the central government budget cover the rest. In 2012 that transfer amounted to 1,416.2 million EUR, or 29.2% of total PAYG revenues (Ministry of Finance, 2013).

3.2. Taxation of Pensions

Contributions for the PAYG pillar are entirely deductable from the personal income tax base, while pensions from the PAYG pillar are subject to personal income tax under

[13] For 40 years of work women receive total accrual rate of 64.25% if they retire in 2013-2016 period, 63.5% in 2017-2019 and 61.5% for retiring in 2020-2022 period. Nevertheless, minimum and maximum pension base as of December 2012 are 551.2 EUR and 2,204.4 EUR, respectively. Taking into account that there is no ceiling for the PAYG contributions, such a pension base setting mechanism has a strong redistributive effect.

[14] Exemptions are some selected professions, such as miner, or soldiers, where additional compulsory contributions paid by employers' are collected by special government owned pension fund.

[15] The self-employed on the other hand pays the same rate of contributions from the base which is a function of annual income from self-employment with the ceiling equal to 2.4 average national gross wage.

an advantageous tax-credit system. As a result, most pensions (approximately 97%) are effectively tax-free, whereas the remaining 3% are taxed at a relatively low effective tax rate. On the other hand, the contributions to the second pillar are deductable from the personal income tax base up to the certain level. This tax relief is limited with the 5.844% of employee's annual gross wage or the nominal amount set annually (2,755.7 EUR in 2013) – whatever it is lower. Pensions from the second pillar are not entitled to the same tax credit as pensions from the first pillar. Instead, 50% of the second-pillar pension is subject to tax, without any special tax credit. As a result, these pensions are taxed more than the first-pillar pensions.

Table 2 includes average gross and net wages in 2013, as well as the maximum amount of tax relief for the second-pillar contribution. One can observe that only taxpayers from the highest decile group can take full advantage of the nominal tax relief for second-pillar contributions.

Table 2: *Gross average annual wage, net average annual wage and maximum amount of tax relief for the second-pillar contribution in 2013 (in EUR)*

Decile group	Average gross wage	Average net wage	Maximum tax relief (5,844 %)
1	9,671.5	6,455.4	565.2
2	10,760.0	7,140.1	628.8
3	11,939.4	7,875.8	697.7
4	13,190.8	8,642.1	770.9
5	14,563.9	9,445.3	851.1
6	16,288.2	10,417.1	951.9
7	18,546.8	11,661.9	1,083.9
8	21,740.3	13,385.6	1,270.5
9	26,755.5	15,872.7	1,563.6
10	47,663.8	24,743.2	2,785.5

Source: Authors' calculation based on data from Statistical Office (2013).

4. THE IMPACT OF DEMOGRAPHIC CHANGES ON BENEFITS FROM THE PAYG PILLAR

The twentieth century experienced explosive population growth, but the twenty-first century is likely to see the end of population growth and instead population aging (Lutz et al., 2004). According to population projections, in the future there will be strong demographic pressures on public expenditures for pensions, health care, and long-term care (European Commission, 2012). Scholars began warning of this decades ago, but we have seen no changes, mainly because short-term-oriented politicians have as their horizon only the next elections. They are not interested in projections for a distant future.

The situation, though, has become so aggravated that taking action cannot be further postponed. Many countries have already taken various measures. International organizations are pressuring countries to act in a timely manner to facilitate and accelerate change.

PAYG systems are vulnerable to population aging. In our analysis, we apply Eurostat population projections from EUROPOP2010 for 2010–2060. They were prepared by the Eurostat for the European countries (EU27) and European Free Trade Association countries (EFTA)[16]. The projections assumed gradual convergence of countries' mortality and fertility, with the year 2150 set as the convergence year. However, the projections were prepared only until 2060, when only partial convergence has been reached.

In Slovenia the life expectancy at birth is increasing rapidly. The past decade alone (from 2000–2001 to 2011) saw an increase of almost 4.5 years for males (72.1 to 76.6 years) and 3.3 years for females (from 79.6 to 82.9 years) (Statistical Office of the Republik of Slovenia, 2012, p. 79). Some developed countries already have a considerably higher, and still-increasing, life expectancy than Slovenia.[17]

The current population age structure is given. The baby-boom generations, born after World War II during times of high fertility, are now in their 50s and early 60s. Over the coming decade, they will be intensively entering retirement. At the same time, people born during the 1980s are starting to enter the labour market. During the 1980s and 1990s, fertility declined; in the first half of the 2000s, it stabilized at very low levels. In 1980 total fertility rate (TFR, or the average number of children a woman gives birth to, during her fertility period) was 2.1, which was still a replacement-level fertility. Since then, TFR declined until 2003, when it reached only 1.2 (Statistical Office of the Republic of Slovenia, 2008, p. 56). Consequently, the number of newborns decreased sharply in that period. In 2003 just 17,321 children were born in Slovenia, whereas the figure was 30,604 in 1979 (Statistical Office of the Republik of Slovenia, 2012, p. 78). Those reduced generations (they are only about one half of their parent's generations) will also determine fertility levels in the coming two to three decades. Even if fertility (TFR) were to increase, which the projections assume, the absolute number of newborns is expected to fall considerably because there will be fewer women of reproductive age.

Sensitivity analysis (in which we variate fertility assumptions while keeping other assumptions unchanged) shows that, despite the impact of fertility on population size in the long run, we cannot expect increased fertility to considerably mitigate the process of population aging in Slovenia in the coming decades (Sambt, 2008). Further, from an economic point of view, increased fertility does not have positive economic

[16] Iceland, Liechtenstein, Norway and Switzerland.

[17] E.g., in Japan the life expectancy at birth in 2009 was 79.6 years for males and 86.4 years for females (OECD, 2011).

effects for about 20 years, as cohorts of newborns start to enter the labour market. In the meantime, the economic effect can even be negative, causing higher public expenditures in the form of education and other transfers like child allowances and health care.

Immigration decreases the overall aging of the population, especially because most immigrants are relatively young (Eurostat, 2011). However, without assuming unreasonable high immigration, the positive effect is only moderate. With time, immigrants are also aging and entering the age group of 65 and older (Bonin et al., 2000).

Figure 2 presents the projected dynamics of the age structure of the Slovenian population by three broad age groups related to economic activity:[18] 0–19, 20–64, and 65 and older. According to EUROPOP2010 projections by 2060 the Slovenian population should slightly increase – by 11,000 people, which is 0.5% of the total population. However, the change in the age structure of the population is striking. The percentage of people age 65 years and older is expected to almost double in the 2012–2060 period, from 16.6% to 31.6%. In contrast, the size of the working-age population (age 20–64) is expected to shrink considerably – from 64.3% in 2012 to 49.8% in 2060. The combination of those two processes will have serious consequences for the long-term sustainability of public finance systems, unless adjusted accordingly. Sensitivity analysis reveals that those results are very robust for a broad range of assumptions about fertility, mortality and migrations since they are mainly driven by the increasing longevity, and especially by the given population structure (Sambt, 2008, Sambt, 2009).

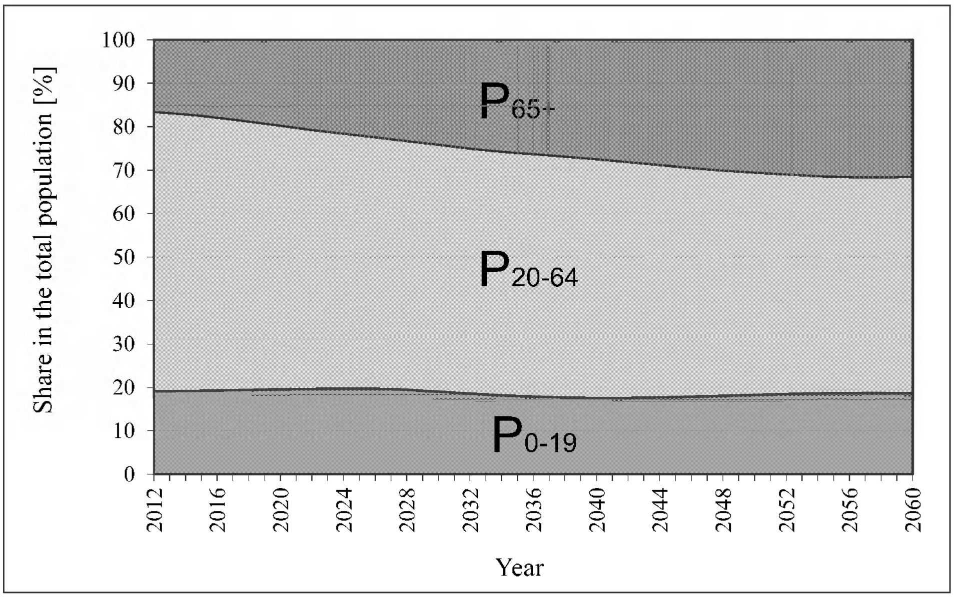
The unfavourable economic development with respect to the population age structure can be shown with the old-age dependency ratio, which is calculated as the ratio between the elderly (age 65+) and the working-age population (age 20–64), multiplied by 100.

An increasing old-age dependency ratio indicates an increasing demographic burden on the productive part of the population in order to maintain the pensions of the economically dependent. According to the EUROPOP2010 population projections, the old-age dependency ratio in Slovenia will increase from 25.9 in 2012 to 63.4 in 2060 (see Table 3).

A rapidly increasing old-age dependency ratio is not specific only to Slovenia. Practically all developed countries across the globe face strong population aging. Therefore, the analysis we present here is generalizable. Table 5 presents projected future increases in the old-age dependency ratio for all EU27 member states, including Slovenia. In the new EU member states (EU12), the increase is expected to be somewhat stronger than in the old EU member states (EU15).

[18] In demography traditionally defined dependency ratio compares population aged 65+ with population aged 15-64. However, in developed countries using 20-64 years in the denominator has been seen as more adequate from the economic point of view since not many individuals enter the labour market before age 20.

Figure 2: *Slovenian population in broad age groups: EUROPOP2010 projections for 2012–2060 (%)*



Source: Eurostat, 2011.

Table 3: *Old-age dependency ratio in EU countries: EUROPOP2010 projections for 2012, 2030 and 2060*

Old member states (EU15)				New member states (EU12)			
	2012	2030	2060		2012	2030	2060
Belgium	29.2	40.5	48.5	Bulgaria	28.7	42.4	66.3
Denmark	29.6	40.7	48.0	Czech Republic	25.0	37.6	60.4
Germany	33.8	51.0	65.1	Estonia	27.6	39.7	61.4
Spain	27.6	38.6	61.6	Greece	31.9	41.9	62.2
France	29.4	43.4	51.7	Latvia	27.4	39.4	74.3
Italy	34.0	44.5	64.5	Lithuania	26.0	38.8	62.3
Cyprus	21.7	33.7	48.7	Hungary	26.8	36.4	62.8
Ireland	19.9	31.2	48.5	Malta	25.8	42.8	60.5
Luxembourg	22.5	32.7	43.0	Poland	21.3	38.6	70.5
Netherlands	26.8	44.2	51.7	Romania	23.3	32.8	70.4
Austria	28.7	42.1	55.4	Slovenia	25.9	42.5	63.4
Portugal	29.9	40.9	59.7	Slovakia	19.0	34.2	67.4
Finland	30.6	47.3	54.5				
Sweden	32.3	41.3	51.7				
UK	28.6	38.6	46.5				

Source: Eurostat, 2011.

4.1. Projecting Future Public Pension Expenditures

Strong population aging translates into pressure on the public pension system. The model that we use in the simulations rests on the age profiles from the base year. Therefore, we refer to it as the age-profiles-based model. Such models are standard approach in Generational accounting methodology[19]. This model is also used for projecting pension expenditures for Slovenia published in 'The Ageing reports' by the European Commission (Ageing Working Group).

In calculations we use three types of matrices. The matrix of pension age profiles (PENS) includes average pensions by years in the future. It builds on the pension age profile from the base year (2011). In particular, the *PENS* matrix consists of two matrices multiplied with each other. The first one contains age profiles of average pension benefits per receiver, whereas the second one includes the share of pensioners in the total population by age group (i.e., retirement rates). This decomposition enables us to more easily, and more accurately, introduce future changes into the age profiles (e.g., increasing retirement age).

Every year those age profiles are shifted up by the assumed growth of pensions. Before 2013 all pensioners with the same retirement conditions received the same level of pension, regardless of the time of retirement ('horizontal equalization'), which strongly simplified the calculations. Horizontal equalization in Slovenian pension system was achieved through complex mechanism of valorisation that was abolished in 2013. Now growth of pension is in real terms[20] 60% indexed to the growth of wages. We follow each cohort of pensioners separately. We use the standard macroeconomic assumption that wages grow in line with labour productivity growth – we use the latest European Commission assumptions.

The coefficient matrix (C) summarizes the effects of future departures from the basic age profile, assumed in the pension matrix. It contains the impact of the Slovenian pension system on pension age profiles in the future. The legally enforced, but gradually phasing-in parameters of the Slovenian pension reform are a typical such case. We obtained several inputs for the coefficient matrix (C) from simulations on microdata on pensioners who have already retired. We simulate their behaviour under pension parameters that will be valid in future years. Weighted averages of those results (by age groups) enter the coefficient matrix.

The population matrix (P) contains the EUROPOP2010 population projections presented earlier.

[19] For review of Generational accounting methodology see, for example, the initial paper from Aurebach, Gokhale and Kotlikoff (1991) and comparative studies across countries (European Commission, 2000; Auerbach, Kotlikoff, & Leibfritz, 1999).

[20] Formally, the growth of pensions is 60% indexed to the nominal growth of wages and 40% to the growth of consumer price index (inflation). This is equivalent to 60% indexation in real terms.

The amount of pension expenditures on individuals aged *k* in year *t* is thus calculated as follows (matrices are multiplied in an element-by-element manner):

$$PENSEXP_{a,t} = PENS_{a,t} P_{a,t} C_{a,t} G_t, \tag{1}$$

where G_t contains coefficients of the cumulative growth of a pensions from the base year (here, 2011) to time *t*. Public pension expenditures (PENSEXP) in year *t* are calculated as the sum of projected expenditures (revenues) by all age groups:

$$PENSEXP_t = \sum_{a=0}^{D} PENSEXP_{a,t} \tag{2}$$

where index *a* runs from 0 to *D*, and *D* denotes the maximum length of life (in our model, the age group 100 and older).

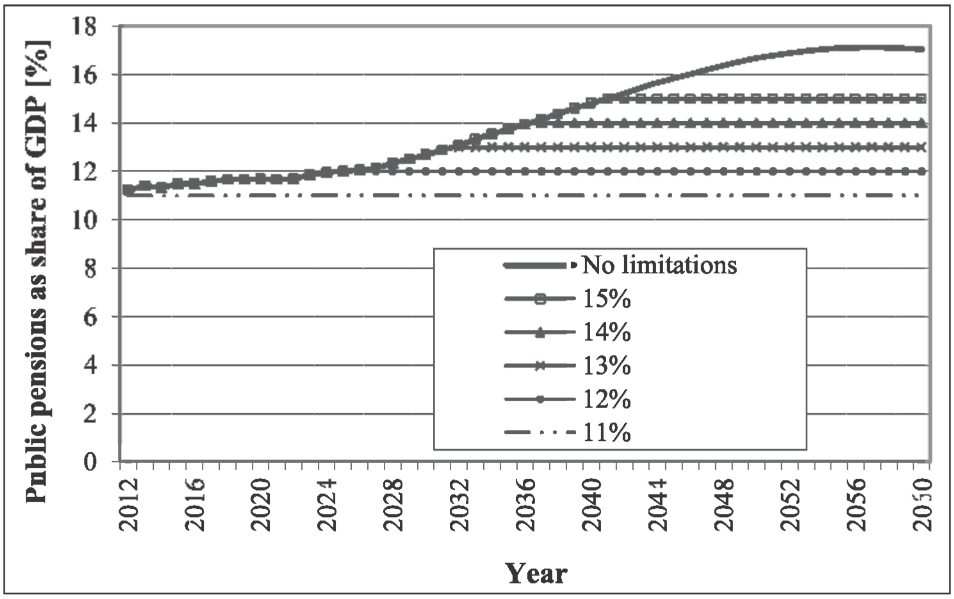
Finally, we employed the set of macroeconomic assumptions from the European Commission prepared in 2011, including assumptions on productivity growth, GDP, employment, and unemployment rates[21].

In the future the age profile of employment rates is projected to shift into higher ages. Consequently, also the age profile of retirement rates will withdraw into higher ages. Both of those two effects are entering the model through *PENS* matrix. Increasing employment rates (also because of assumed decline in unemployment rates) will have positive impact also on GDP since GDP growth in the model consists of labour productivity growth and the employment growth – the same approach uses also the European Commission in their calculations. For example, in the transition period up to 2018 the level of pensions will gradually decline relative to wages because of gradually increasing number of years that are taken into account when calculating pension base (from 19 best years in 2013 to 24 in 2018). This effect enters the model through *C* matrix as well as other effects of the pension legislation that were calculated using the microdata on pensioners. Nonetheless, it is projected that the public pension expenditures as a percentage of GDP will strongly increase since the results are mainly driven by the population ageing.

In 'pension' category we include old-age, disability and survivor pensions but also pensions of farmers, police officers, World War II veterans, state pensioners and so on. However, payments to the health insurance that are paid for the pensioners by the pension fund are not included. Using the age-profiles-based model, we obtained the results presented in Figure 3, labelled 'No limitation' variant. Without further changes to the current pension system, population aging would largely translate to rapidly growing public pension expenditures as a % of GDP.

[21] To link employment rates with retirement rates, we used the submodel of the Institute of Macroeconomic Analysis and Development (IMAD; for a detailed description of the submodel, see Kraigher (2005).

Figure 3: *Projections of public pension expenditures in Slovenia in 2012–2060 (% of GDP), scenario* with *no limitations and scenarios if limiting public pension expenditures to certain % of GDP*



Source: authors' calculations.

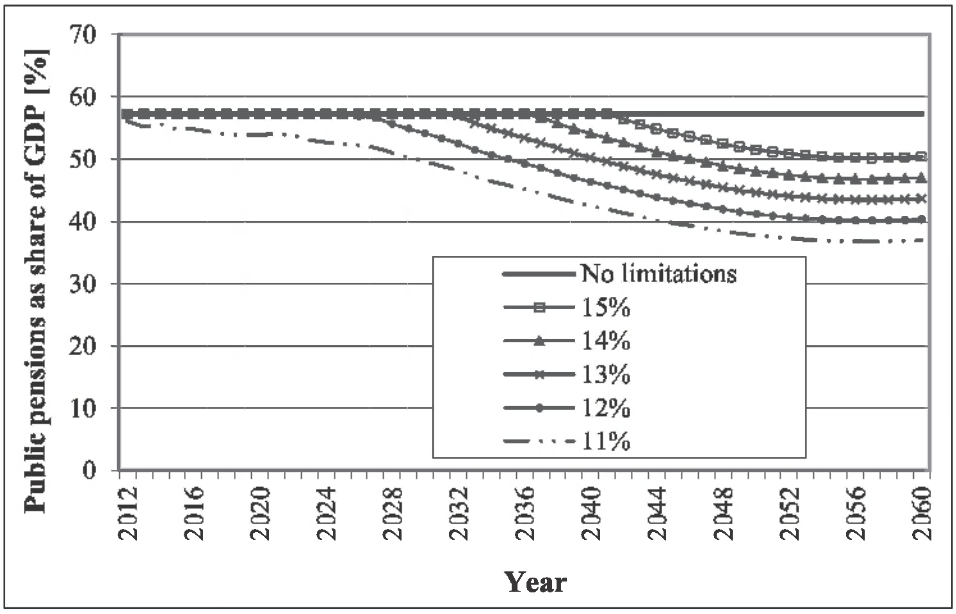
There are basically three solutions for mitigating rapidly growing pension expenditures. The first, usually considered preferable, is to increase the retirement age. This solution also provides the most straightforward response to increasing longevity. The second solution is to increase taxes, usually on labour income. In Slovenia labour is already highly taxed, which hinders its international competitiveness. Moreover, the tax burden has a negative impact on employment and incentives to work. The third solution is to reduce the level of pension benefits from the PAYG system.

In our analysis we focus on the third option by introducing a simple assumption about future reductions of pension benefits. We assume the government will have to prevent further increases in public pension expenditure above some percentage of the GDP (i.e., capping expenditures) in a way that all pensions will be cut proportionally, regardless of the type and level of pension. Thus, we set the tolerated maximum percentage of public pensions of GDP at 11%, 12%, 13%, 14%, and 15% of GDP. Figure 3 shows projected public pension expenditures as a percentage of GDP according to these scenarios.

4.2. Expected Level of Pensions from the PAYG Pillar

As already explained in section 3 in 2013 the statutory accrual rate for a man with 40 working years was set to 57.25% and according to the current pension law it will remain at this level in the future ('No limitations' scenario in Figure 4). We also present results for different scenarios of assumed maximums to which government would tolerate pension expenditures to increase; in those cases, the net replacement rate would fall to even (much) lower levels, as it is revealed in Figure 4. In the case that expenditures for pensions would be capped at 11% of GDP, the net replacement would thus decline to 37% by 2060.

Figure 4: *Projections of net replacement rate at retirement in the period 2012–2060*



Source: authors' calculations.

The presented PAYG-based net replacement rates are very low, and individuals without other means will at best not be able to sustain their living standards. Many of them are expected to fall below poverty threshold. Achieving the 70% net replacement rate suggested by the OECD will be possible only with regular private pension savings, which should fill the PAYG shortfall. Before we present the analysis about the required saving during the working period, we first present the characteristics of traditional asset classes, as they have an impact on amounts of savings needed.

5. CHARACTERISTICS OF TRADITIONAL ASSET CLASSES OVER THE LONG RUN

In this section we analyze three traditional asset classes (i.e. stocks, treasury bonds, and treasury bills). The purpose is not to simulate optimal asset allocation over the long run but to show the impact of the asset allocation decision stemming from the characteristics of the previously mentioned asset classes.

We base our approach on historical yields (arithmetic and geometric) and volatilities reported in the literature and build the yield and volatility model. We use different global historical datasets as a source[22]: US data for the period 1802–2001 and the period 1946–2011 (Siegel, 2002), US large-cap and world data for the period 1926–2005 (Bodie, Kane, and Marcus, 2009), US and world data for the period 1900–2000 (Dimson, Marsh, and Staunton, 2002), US large-cap data for the period 1926–2005 (Malkiel, 2007), and MSCI stock indices for the period 1969–2010. We calculated two to 40-year yields and standard deviations using every data source and then averaged yields and standard deviations. Yields are calculated according to the fact that they should fall over time, as over time the geometric average yield (which is lower than arithmetical average yield) becomes more realistic than the arithmetic average. We borrow the formula from Bodie, Kane, and Marcus (2009).

We calculate standard deviation according to the random-walk assumption.[23] In the short run stocks are more volatile than the other two asset classes, which calls for a higher required yield: yield together with dividends (representing one-third of total nominal return) historically has been around 10%. Over shorter horizons (even 10 years), investment performance can be quite different (e.g. MSCI US Standard Core Total Return index returned -1.29 in nominal terms on average between 31.12.1999 to 31.12.2009, but also as high as 19% between 31.12.1989 and 31.12.1999.[24]

Figure 5 shows that standard deviation is not persistent if we consider longer investment horizons. Namely, in 15 years, yield distribution has approximately one-fourth of one-year standard deviations. Thus, in the longer run, the changed relationship between yield and risk of stocks relative to bonds or bills favours stock.[25] Siegel (2002) argues that empirically verified long-term standard deviation is much lower that standard deviation assumed by the random-walk model, and that after slightly less than 20 years, standard deviation of stocks even falls bellow standard deviation of bonds.

We have deliberately chosen the conservative *i.i.d.* assumption and used a 6.53% expected average real yield for 20-year investment in stocks, 1.25% for 20-year investment in T-bonds, and 1.11% for 20-year investment in T-bills. Over the 40-year investment horizon, yields used were 6.17%, 1.17%, and 1.07% respectively (see Table 6). All the yields are expressed net of management fees, which we assumed to be 1.3% for stocks, 1.0% for T-bonds, and 0.5% for T-bills. After calculating average yields, we calculated standard deviation and then minus-one and minus-two standard deviation yields (–1 sigma and –2 sigma yields) for each asset class for various investment horizons (see Table 4).

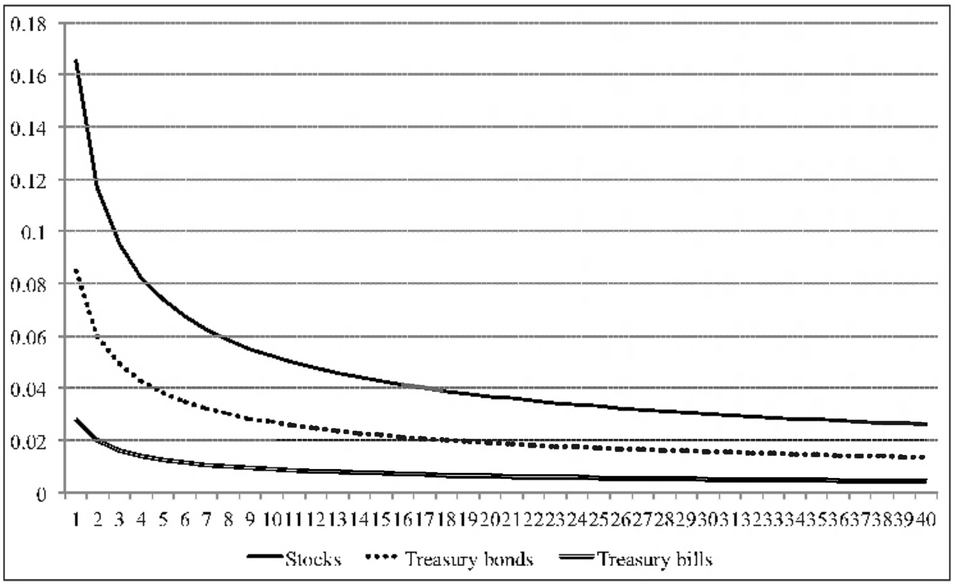
[22] Slovenian pension funds' investment policies should be global and diversified. Therefore, global historical data are the most reasonable data input in our analysis.

[23] I.e., as a square root of forecasting period multiplied by one-year standard deviation of used indices, as distribution of returns are assumed to be - identically independently distributed (*i.i.d.*).

[24] Calculations using MSCI indices are not shown for brevity.

[25] There are still multiple differences in yields but differences in standard deviations become much smaller.

Figure 5: *Standard deviation of traditional asset classes over longer periods*



Source: Malkiel (2007), world markets data, and authors' calculations.

Table 4: *Real yields and volatilities of traditional asset classes during investment horizons of 20 and 40 years (%)*

		Large- and mid-cap stocks	T-bonds	T-bills
	Average yield	6.53	1.25	1.11
20 years	–1 sigma yield	2.39	-0.67	0.43
	–2 sigma yield	-0.84	-2.59	-0.26
	Average yield	6.17	1.17	1.07
40 years	–1 sigma yield	3.24	-0.19	0.59
	–2 sigma yield	0.31	-1.55	0.10

Note: 20-year yields are somewhat higher than 40-year yield, as in the shorter run there is higher effect of arithmetic average yield, which is always higher than geometric average yield (Bodie, Kane, and Marcus, 2009). Figures are net of management fees. For stocks, we assumed annual management fees to be 130 basis points (bps); for T-bond portfolios, 100 bps; and for T-bill portfolios, 50 bps.
Sources: Siegel (2002); Bodie, Kane, and Marcus (2009); Dimson, Marsh, and Staunton (2002); Malkiel (2007); authors' calculation.

6. THE MODEL

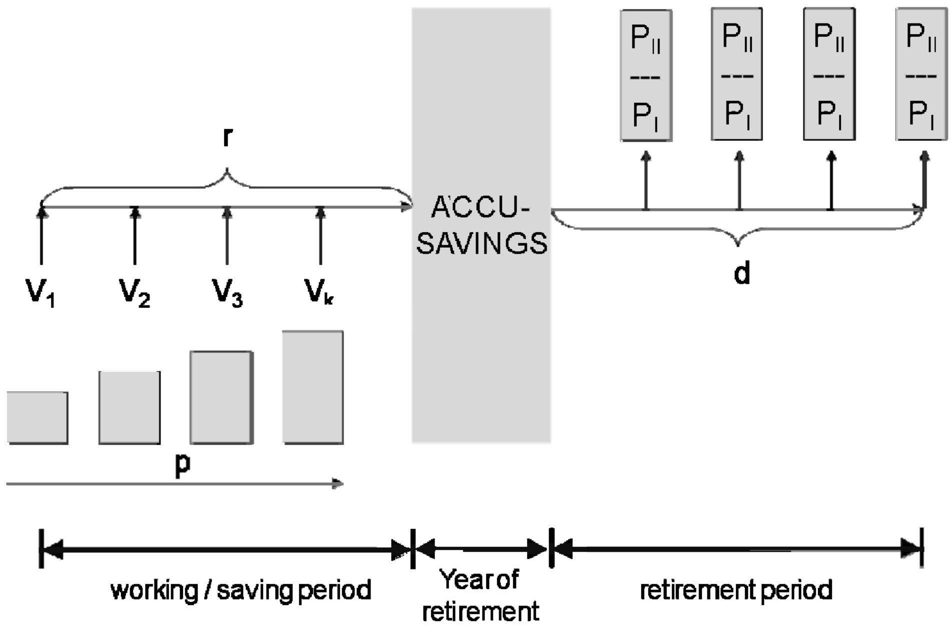
Figure 6 shows the basic model we started with. Taking into account the unsustainability of the PAYG pillar, we expect that future pension recipients will receive pensions from the PAYG pillar substantially lower that the 70% net replacement rate recommended by the OECD. Therefore, we estimate the monthly pension gap, PGAPt (i.e., the difference between the 70% net replacement rate and the forecasted PAYG replacement rate) for a typical male pension beneficiary and assume that he is motivated to increase his periodic pension savings over the entire working period to a level sufficient to cover the future pension gap:[26]

$$PGAP_t = (Pension_t^{70\%} - Pension_t^{PAYG})$$

Apart from gender and retirement age, which affect the pension gap (PGAPt), we also take into account three different public finance scenarios that affect the individual PAYG monthly pensions differently. Because of the fiscal unsustainability issues addressed in section 4, we decided to work with three hypothetical public finance scenarios:

- The no-limit scenario assumes no limits for the percentage of PAYG pension expenditures as a percentage of GDP for the future.
- The 13% GDP scenario assumes the proportion of PAYG pension expenditures to be capped at 13% of GDP.
- The 11% GDP scenario assumes the proportion of PAYG pension spending to be capped at 10% of GDP.

Figure 6: *A graphical illustration of the model*



Legend: r = investment rate of return; p = the growth rate of salaries and saving instalments; V_i = monthly saving instalments; d = discount rate; P_I = monthly PAYG pension; P_{II} = pension gap.

[26] Calculations do not conceptually differ for female individuals, but we excluded those results from this paper for sake of keeping results as short and concise as possible.

Of course, by increasing restrictions on the pension-to-GDP ratio, the actual forecasted net replacement rate deteriorates gradually and the monthly pension gap increases accordingly. Consequently, the additional pension savings that have to be accumulated through the private pension system must increase by increasing the level of public finance restrictions, assuming that individuals target their individual total pensions at the 70% net replacement rate.

In the next step the monthly pension gap values ($PGAP_t$) for individual pension recipients are discounted using a 0.5% technical discount rate[27] to the total amount of savings needed at the year of retirement (ACCUSAVINGS), which represents the target value an individual must accumulate over his working period through his monthly savings in the private pension pillar. In discounting, we use male life expectations at the age of 61 according to the Deutsche Aktuarvereinigung (DAV) tables. Again, we take into account the effect of the length of the savings period for typical individuals, and we simulate investment strategies that are consistent with three different asset allocation strategies. The simulated investment strategies rely on three asset classes, characterised by their distinct risk-return profiles: (1) stock strategy, (2) bond strategy, and (3) bill strategy. For simplicity, investors are assumed to stick to the selected asset class (i.e., risk-return profile) throughout the entire investment horizon, and they are assumed not to mix the three asset classes.

In our results we present the amounts that must be saved by a male individual in the private pension system for a 20-year and a 40-year working period. We assume the starting annuity (A_t=1) to grow monthly by the expected average growth rate of salaries (g), which should be in line with productivity growth rate (we assume the average salary grows by 1.77% per year), and we assume those annuities to be invested at the constant investment rate r, which depends on the preselected asset class and related risk-return profile. Table 6 presents yields and volatilities for selected asset classes for selected 20- and 40-year investment horizons:

$$A_{t-1} = \frac{ACCUSAVINGS * (r-g)}{(1+r)^n - (1+g)^n}$$

Table 5 displays a summary of the results. The results are presented for male individuals in three different decile groups (D1, D5, and D10) for selected years in the period 2035–2060. Evidently, the pension gap ($PGAP_t$) is inflated throughout the projection period for all income groups (D1, D5, and D10), as the net replacement rate from the PAYG system is projected to deteriorate. In nominal terms the gap is becoming larger for individuals who belong to higher-income groups. This means that unless such individuals accumulate greater savings until the end of their working period they will fall below the 70% net replacement rate.

As previously explained, the discounted pension gaps represent the accumulated savings that each individual pension recipient is expected to accumulate during his working

[27] We use 0.5% discount rate as it reflect the need to minimize risk exposure once the individual is retired and it is consistent with annuity industry practice.

period until the end of his retirement year. Consequently, the volume of the required accumulated savings determines the monthly savings contributions each male individual is expected to save until the retirement year. Table 8 presents the first annuity a male individual is expected to start saving under varying assumptions. First, we assume that individuals from different decile groups have to accumulate different savings volumes to supplement the regularly expected PAYG pension. Second, we assume that future public finance scenarios affect the monthly savings contributions. And third, the length of the expected savings period also affects the volume of the accumulated funds at the end of the working period. For simplicity, we present only calculations for 20 years and 40 years.

Table 5: *PAYG pensions calculated by the official net replacement rate in selected years and gaps to the 2035 pension, 70% net replacement rate pension, and gap to the forecasted salary in selected years to 2060 (in EUR)*

		2035	2040	2045	2050	2055	2060
Decile group	Net replacement rate (%)	57.25	55.29	53.39	51.56	49.79	48.08
D 1	PAYG pension (M)	461	486	513	540	570	601
	Gap to the 2035 pension	0	25	51	79	109	139
	Gap to the 70% net replacement rate	103	129	159	193	231	274
	Gap to the salary	344	393	448	508	575	649
	Salary	806	879	960	1048	1144	1249
D 5	PAYG pension (M)	675	711	750	791	834	879
	Gap to the 2035 pension	0	37	75	116	159	204
	Gap to the 70% net replacement rate	150	189	233	283	338	401
	Gap to the salary	504	575	655	743	841	949
	Salary	1179	1287	1405	1534	1674	1828
D 10	PAYG pension (M)	1768	1864	1965	2071	2184	2302
	Gap to the 2035 pension	0	96	197	304	416	535
	Gap to the 70% net replacement rate	394	496	611	741	887	1050
	Gap to the salary	1320	1507	1715	1946	2203	2486
	Salary	3088	3371	3680	4018	4386	4789

Source: Authors' calculations.

As it is evident from Table 8, the individual's decision for a particular type of investment (i.e., asset class) and the length of the savings period have a substantial impact on the size of the annuity that the individual saver is expected to start saving. So, a male individual in the D5 decile group who decides to invest in a portfolio of large- and mid-cap stocks (see section 5) is expected to start saving 54 EUR per month if he has a 40-year investment period and 121 EUR per month if he has a 20-year investment period. If the same male individual were to decide to invest in a portfolio consisting exclusively of T-bills, he would need to start saving 149 EUR with an intended investment period of 40 years and 222 EUR per month with an intended investment period of 20 years. The differences in required monthly savings contributions are significant, and one can clearly observe how important it is to decide on a proper investment strategy in terms of both portfolio structure and length of the savings period (i.e., individuals should start saving as soon as possible). All other accompanying aspects that also affect the final savings outcome (e.g., different public finance scenarios that directly affect the PAYG pensions) make the differences only more pronounced.

The second set of results is based on simulations in which the investment yields were adjusted to reflect the volatility of average historical returns of the preselected asset classes. Therefore, the three right-hand columns of Table 6 present the required monthly savings contributions for a risk-aware male individual who wants to avoid a case that investment yield deviates down to two standard deviations (–2 sigma) from the average historical returns of the individual asset classes. In this scenario all required monthly savings contributions are significantly higher, which reflects the sensitivity of the savings strategy to financial market volatility.

Table 6: *Required contributions under three different fiscal scenarios consistent with average real yield under three different asset class allocations (left) and consistent with –2 sigma real yield under three asset class allocations (right)*

		Average yields			–2 sigma yields		
		D1	D5	D10	D1	D5	D10
40 years	SC I - STOCKS - AVERAGE						
	1st contrib. under "no limit"	37	54	142	116	169	444
	1st contrib. under 13% GDP	57	84	219	179	262	685
	1st contrib. under 11% GDP	71	104	271	221	324	848
	SC2 - BONDS - AVERAGE						
	1st contrib. under "no limit"	100	147	384	154	226	591
	1st contrib. under 13% GDP	155	226	593	238	348	913
	1st contrib. under 11% GDP	191	280	734	295	431	1130
	SC3 - BILLS - AVERAGE						
	1st contrib. under "no limit"	102	149	391	120	175	459
	1st contrib. under 13% GDP	157	230	603	185	271	709
	1st contrib. under 11% GDP	195	285	746	229	335	877
20 years	SC1 - STOCKS – AVERAGE						
	1st contrib. under "no limit"	83	121	318	184	269	705
	1st contrib. under 13% GDP	128	187	490	284	416	1089
	1st contrib. under 11% GDP	158	232	607	352	514	1347
	SC2 - BONDS - AVERAGE						
	1st contrib. under "no limit"	150	219	574	184	269	705
	1st contrib. under 13% GDP	231	338	886	284	416	1089
	1st contrib. under 11% GDP	286	418	1096	352	514	1347
	SC3 - BILLS - AVERAGE						
	1st contrib. under "no limit"	152	222	582	174	255	667
	1st contrib. under 13% GDP	234	343	898	269	393	1030
	1st contrib. under 11% GDP	290	424	1112	332	486	1274

Note: D1, D5 and D10 represent first, fifth and tenth decile group of an individual's income distribution.
Source: authors' calculations.

Finally, we compare accumulated savings (i.e., pension wealth), assuming that an individual would start saving monthly contributions adequate to the expected extreme market performance (i.e., –2 sigma) and at the same time it would *ex post* turn our that he could realise expected average market yields (mean yields). The results shown Table 7 are striking. Under this approach, one can easily grasp the advantages of allocating pension wealth into stocks over the long run.

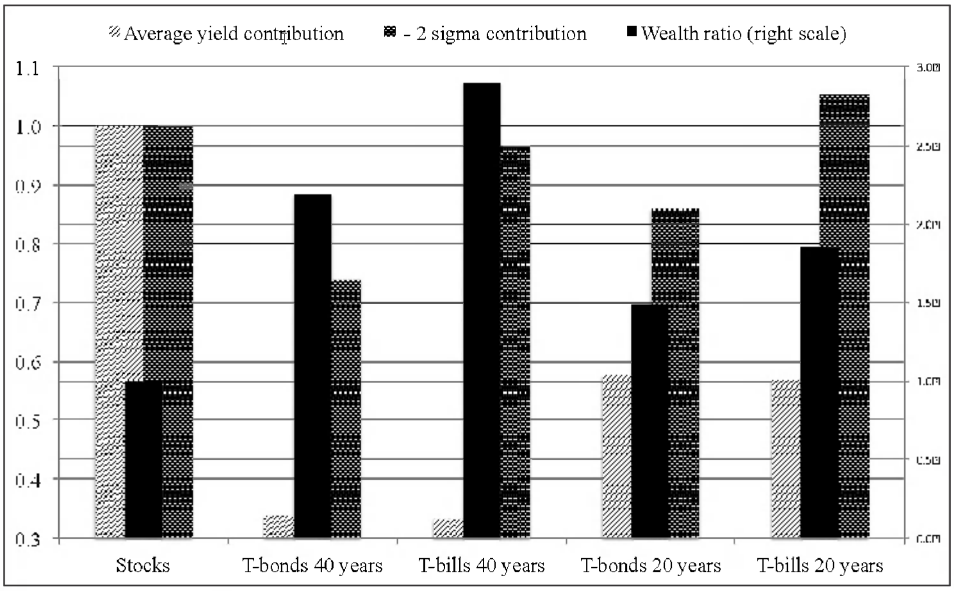
Table 7: *Pension wealth at the moment of annuitizing of an individual who anticipated extreme market performance (–2 sigma real yield) but realized average performance under three different asset class allocations and three different fiscal scenarios*

		D1	D5	D10
40 years	SC1 - STOCKS - (–2 sigma)			
	1st contrib. under "no limit"	330,880	484,132	1,268,246
	1st contrib. under 13% GDP	510,912	747,547	1,958,297
	1st contrib. under 11% GDP	632,217	925,036	2,423,253
	SC2 - BONDS - (–2 sigma)			
	1st contrib. under "no limit"	149,147	218,226	571,672
	1st contrib. under 13% GDP	230,297	336,963	882,718
	1st contrib. under 11% GDP	284,977	416,967	1,092,300
	SC3 - BILLS - (–2 sigma)			
	1st contrib. under "no limit"	113,739	166,419	435,955
	1st contrib. under 13% GDP	175,624	256,967	673,158
	1st contrib. under 11% GDP	217,322	317,978	832,985
20 years	SC1 - STOCKS - (–2 sigma)			
	1st contrib. under "no limit"	109,335	159,975	419,076
	1st contrib. under 13% GDP	168,824	247,018	647,095
	1st contrib. under 11% GDP	208,908	305,667	800,734
	SC2 - BONDS - (–2 sigma)			
	1st contrib. under "no limit"	73,977	108,241	283,551
	1st contrib. under 13% GDP	114,228	167,135	437,831
	1st contrib. under 11% GDP	141,349	206,817	541,785
	SC3 - BILLS - (–2 sigma)			
	1st contrib. under "no limit"	58,765	85,983	225,245
	1st contrib. under 13% GDP	90,740	132,767	347,800
	1st contrib. under 11% GDP	112,284	164,289	430,377

Note: D1, D5 and D10 represent first, fifth and tenth decile group of individual's income distribution.
Source: authors' calculations.

Figure 7 summarizes the effects of different investment strategies chosen by male individuals. Stocks are a benchmark in this comparison (i.e., results are expressed in terms of ratio of stocks to other asset classes). First, individuals who choose stocks over a 40-year period are (according to the expected average yield) required to save about one-third the amount of individuals who choose bond or bills. According to expectations of extreme financial market performance, stock investors can still save about one-quarter less (exactly 26% less over a 40-year investment horizon and 14% less over a 20-year horizon). Second, when risk-aware investors decide to save according to expectations that they close the gap despite extreme financial market performance, but those results turn out (most likely) to be average, a stock strategy would beat out a bond and/or bill strategy by a substantial margin. As Figure 8 shows, this margin is already very material at a 20-year investment horizon. Investors with a stock strategy accumulate 48% more pension wealth than those with a T-bond strategy and 86% relative to a T-bill strategy. Over the 40-year investment horizon, the respective differences are substantial: 121% relative to T-bond strategy and 191% relative to T-bill strategy.

Figure 7: *Stock-to-other-asset-class ratios of contributions and of pension wealth before annuitization for investment horizons of 20 years and 40 years*



Source: authors' calculations.

We argue that governments in countries with pension system facing similar issues to those of Slovenia should be interested in improving the financial literacy of the public in both aspects, i.e. improving an awareness to save and also knowledge about basic characteristics of financial asset classes. Doing so would prevent opportunity losses in terms of lower available pension wealth and old-age disposable income despite people being aware about the need to save for their pensions.

7. ADVANTAGES OF PRIVATE PENSION SAVINGS ALLOCATIONS OVER ORDINARY SAVINGS ALLOCATION

In this section we present benefits that stem from the fact that separate pension pillars react differently to various economic shocks. We calculate the extent of benefits using common finance literature metrics of a standard deviation of a portfolio. In this approach coefficients of correlation among main assets take the central role. The co-movements that are reflected in the coefficients of correlation reduce the volatility of a combination of assets. We measure and show benefits of diversification that rise from combining exposures to homogeneous assets by the percentage decrease of standard deviation of the resulting portfolio compared to the standard deviation of the asset class alone.

In order to be able to calculate standard deviation of a portfolio we have to define homogeneous assets. Our analysis was based on the traditional financial market asset classes, i.e. stocks (EQ), 10-year government bonds (10yB), money market government bills (MM), and on wages, following approach of Holzmann (2002). We used data from Thomson Reuters Datastream database for France, UK, Germany and The Netherlands, for the period from 1971 to 2011. We calculated annual yields from total return index time series in nominal terms. We used the following stock indices: *CAC 40 Total Return Index* (France), *FTSE 100 Total Return Index* (UK), *DAX 30 Total Return Index* (Germany), *Amsterdam MIDKAP DS Total Return Index* (The Netherlands), Benchmark 10-year Government Total Return Indices, and Benchmark 1-3-year Government Total Return Indices. For the time series of wages (we used two wage time series: *Wage Rate – Private Sector* and *Wages and Salaries – Total Economy*) we also used time series of aggregate wages in nominal terms.

Our goal is not to define optimal investment strategy for pension portfolio but to show benefits of diversification that rise from the characteristics of each financial asset class by combining it with the pension income that is received from the PAYG. Therefore, we calculate bivariate correlation coefficients of stocks, long-term government bonds, and short-term government bonds with both measures of wage income that proxy for the dynamics of the PAYG. Table 8 reports our results. In general, correlation coefficients are low, which means that benefits of diversification are substantial. We see that correlation coefficients are smallest in the case of combining PAYG with stocks (cross-county average 0.0016), followed by long-term government bonds - 10yB (cross-county average 0.0690) and finally by short-term government bonds – MM (cross-county average 0.2625), which means that benefits of diversification of PAYG pension income are the largest when this income is combined with pension income that derives from stocks investments.

Table 8: *Bivariate correlation coefficients between annual growth rates of aggregate wages and performance of financial asset classes for the period from 1971 to 2011*

	Stocks	10yB	MM	AVERAGE
France	0.2657	0.2327	0.4169	0.3051
WR	0.3184**	0.2584*	0.4914**	0.3561
WS	0.2130	0.2071	0.3424***	0.2542
United Kingdom	0.1931	0.2274	0.1488	0.1898
WR	0.2646*	0.1412	0.1577	0.1878
WS	0.1215	0.3136**	0.1400	0.1917
Germany	-0.2038	-0.1649	0.3268	-0.0140
WR	-0.2826*	-.1693	0.4026***	-0.0178
WS	0.1208	-.1606	0.2510	-0.0101
Netherlands	-0.2486	-0.0194	0.1574	-0.0369
WR	-.1658	-.054	0.2776*	0.0193
WS	0.3313**	0.0153	0.0372	-0.0930
ALL	0.0016	0.0690	0.2625	0.1110
WR	0.0326	0.0441	0.3323	0.1363
WS	-0.0294	0.0938	0.1927	0.0857

Notes: WR – Wage rate of the private sector, WS – Wages & salaries; 10yB – 10-year government bonds, MM – money market government bills; Performance of financial asset classes is expressed in terms of annual changes in total return indices.
*** - significant at 1%, ** - significant at 5%, * - significant at 10%.
Source: Thomson Reuters Datastream, authors' calculations.

In order to be able to express diversification benefits we have to measure standard deviations of separate time series (see Table 9) and then calculate standard deviations of portfolios comprising PAYG pension income and pension income that derives from a particular financial asset class.

Table 9: *Standard deviations of annual aggregate wages and performance of financial asset classes for the period from 1971 to 2011*

	Stocks	10yB	MM	WR	WS
France	0.2574	0.0845	0.0341	0.0487	0.0524
United Kingdom	0.1700	0.1120	0.0341	0.0583	0.0873
Germany	0.2512	0.0693	0.0339	0.0309	0.0358
Netherlands	0.3104	0.0792	0.0286	0.0471	0.0424
ALL	0.2472	0.0862	0.0327	0.0463	0.0545

Note: 10yB – 10-year government bonds, MM – money market government bills; Performance of financial asset classes is expressed in terms of annual changes in total return indices.
Source: Thomson Reuters Datastream, authors' calculations.

If we compare portfolio standard deviations with standard deviations of a particular financial asset class, we can directly measure the positive diversification impact of combining two sources of pension income, i.e. decrease of the standard deviation. Table 10 and Figure 8 report such effects for pension income that derives in 50 percent from PAYG and 50 percent from the particular financial asset class.

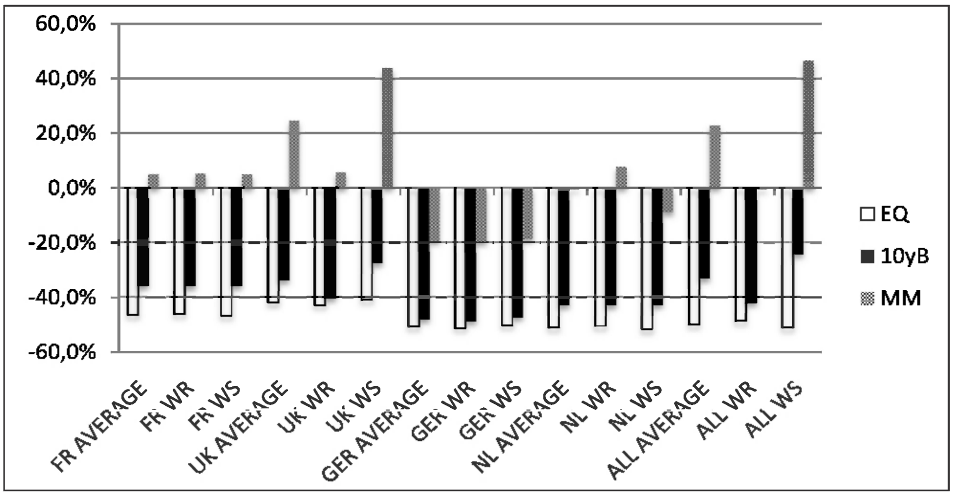
Table 10: *Diversification benefits of combining traditional asset classes (50%) with PAYG (50%)*

	Stocks		10yB		MM	
	STDEVp	% efect	STDEVp	% efect	STDEVp	% efect
FR (0.2574)	0.1376	-46.6%	0.0541	-36.0%	0.0359	5.2%
WR	0.1384	-46.2%	0.0540	-36.2%	0.0360	5.4%
WS	0.1367	-46.9%	0.0541	-35.9%	0.0358	5.0%
UK (0.1700)	0.0985	-42.0%	0.0739	-34.0%	0.0425	24.8%
WR	0.0969	-43.0%	0.0667	-40.5%	0.0360	5.7%
WS	0.1001	-41.1%	0.0811	-27.6%	0.0490	43.9%
GER (0.2512)	0.1234	-50.9%	0.0359	-48.2%	0.0274	-19.3%
WR	0.1220	-51.4%	0.0355	-48.8%	0.0272	-19.9%
WS	0.1247	-50.4%	0.0363	-47.5%	0.0276	-18.8%
NL (0.3104)	0.1513	-51.3%	0.0451	-43.1%	0.0284	-0.6%
WR	0.1531	-50.7%	0.0450	-43.2%	0.0308	7.7%
WS	0.1495	-51.8%	0.0452	-42.9%	0.0260	-9.0%
ALL (0.2472)	0.1235	-50.0%	0.0575	-33.3%	0.0402	23.0%
WR	0.1265	-48.8%	0.0498	-42.2%	0.0325	-0.7%
WS	0.1206	-51.2%	0.0653	-24.3%	0.0479	46.7%

Note: WR – Wage rate of the private sector, WS – Wages & salaries; 10yB – 10-year government bonds, MM – money market government bills; Performance of financial asset classes is expressed in terms of annual changes in total return indices; STDEVp – standard deviation of a portfolio created by combining financial asset class with PAYG; %effect – Percentage decrease of standard deviation of an asset class due to diversification through PAYG.
Source: Thomson Reuters Datastream, authors' calculations.

As expected from the levels of correlation coefficients, decreases of standard deviation and thus benefits of diversification are the largest in case of combining PAYG income with investments into stocks, whereby standard deviation decreases to about one half of standard deviation of stocks. They are followed by investments into long-term government bonds, whereby standard deviation decreases to about two thirds of standard deviation of long-term government bonds. Benefits are limited in case of short-term government bonds, which is also a result of similar levels of standard deviations of wage income and short-term government bond yields. We see from Table 11 that standard deviations of short-term government bonds are generally lower than standard deviations of wage income.

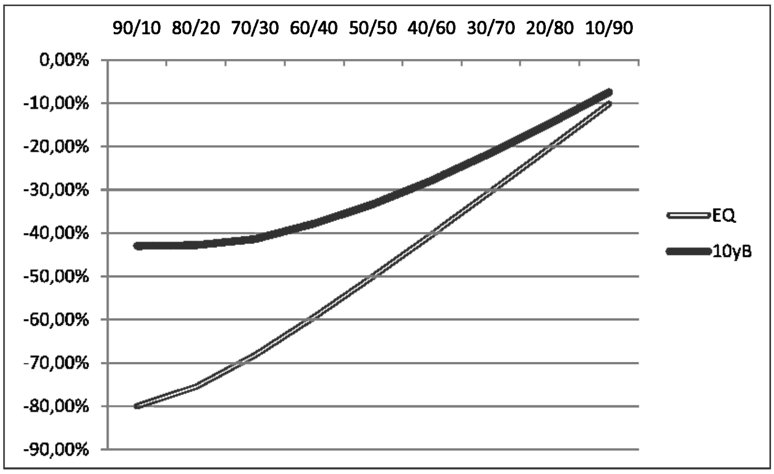
Figure 8: *Diversification benefits of bivariate combinations of financial asset classes (50%) and PAYG (50%), measured as a percentage decrease of standard deviation of a financial asset class*



Note: WR – Wage rate of the private sector, WS – Wages & salaries; 10yB – 10-year government bonds, MM – money market government bills.

However, diversification benefits depend on the weights in the standard deviation equation, i.e. on the relative importance of each source of pension income. They are the highest in situations when country makes the first moves towards private pensions from the PAYG and decrease with the relative importance of private pensions (see Figure 9).

Figure 9: *Impact of the extent of required private pension supplement on diversification benefits*



Note: 10yB – 10-year government bonds; 90/10 – pension income generated in 90% through PAYG and 10% through private pensions, 80/20 – pension income generated in 80% through PAYG and 20% through private pensions, etc..

8. CONCLUSION AND DISCUSSION

With EU demographic dynamics, many countries are expected to face a situation where PAYG system will not be able to finance levels of pensions set out in current rules. In this article, we show that this is the case in Slovenia. Taking into account the current pension system and aging population, PAYG pension benefits as a percentage of GDP would increase to about 17%, which we believe is financially unsustainable. Cuts to the PAYG benefits thus are unavoidable. It is therefore the role of the private pensions to fill the gap between projected first-pillar pension and overall level of pension at the 70% net replacement rate suggested by the OECD. Improved private pension systems would not only solve pension issues but also help develop financial markets, which would in turn lead to higher savings, higher capital budgets of companies, economic growth, and most importantly, well-being of the population.

Today Slovenia's private pension system is of minor importance. Second-pillar legislation was enforced in 2000, a significant step forward, but evolution in the field has since been negligible. New legislation is now effective since January 2013, but secondary level acts are still being prepared. In general, people do not have enough knowledge to make proper decisions. Because a large majority of the population is only modestly financially literate, there is a real risk that many people will not have enough means to live through old age without financial difficulties.

In this paper we have shown how much people should save and what kind of asset allocation they should choose. The government should try to address the issue of financial illiteracy and encourage people to save. In addition, government should conceptualise reasonable legislation on the available financial vehicles offered in the private pension system and work on ways to properly communicate the asset allocation decision. Namely, we have shown that if an individual saves over a period of 40 years and allocates savings into a well-diversified stock portfolio, he can save far more than an individual who allocates savings into a well-diversified T-bond or T-bill portfolio for the same expected horizon. The differences are also significant over a 20-year period.

However, story should not be based solely on returns but should also include risk. Namely, higher stock returns should also be more risky compared to bond and bill returns. That was why we checked the episodes of the worst historic financial market performance. Average annual real yields at two standard deviations below the average value are 0.31%, –1.55%, and 0.10% for stocks, bonds, and bills, respectively. If a risk-aware investor chooses to save amounts consistent with –2 sigma yields (which are higher than amounts consistent with average yield), the required amount for stock allocation is about 26% less than for T-bond allocation over a 40-year investment horizon.

In addition, we have revealed the significant upside in yield potential with stock investing, which is not the case for fixed-income investments. Namely, if risk-aware investor decides to save according to expectations that they save enough to achieve 70% net replacement rate despite extreme financial market performance (but the investment result turns out to be average, which is obviously the most likely), individuals with a stock strategy would beat individuals with a bond and/or bill strategy by a substantial margin. Investors with a stock strategy accumulate 49% more pension wealth than those with a T-bond strategy and 85% more than those with a T-bill strategy over a 20-year investment horizon. Over a 40-year investment horizon, the stock allocation beats the T-bond portfolio by 119% and the T-bill strategy by 190%. We thus conclude that when people who save for their pension have a long-enough horizon, they should predominantly allocate investments into stocks. The amount that people should save every month is in this setting where individuals choose asset allocation for the whole investment horizon determined by asset allocation choice, not income level, which is commonly assumed to determine individual's risk aversion. Governments should bring that finding into legislation, and one way of doing so would be the life-cycle investment policy approach.

According to the results presented here and the characteristics of Slovenia's current pension system, the first thing needed is to give individuals a certain degree of free asset allocation choices and/or life-cycle investment choices for the automatic transition of aging individuals to more conservative asset classes (i.e., toward T-bill allocation). Even though this issue is sensitive, current guarantees set uniformly for all individuals are ill advised. We argue that besides a robust, strong, well-designed second pillar, individual retirement accounts should be introduced in the third pillar. Such accounts, when properly tax-calibrated, would provide lower-income individuals with additional incentives to save for their pension. Over the long run, this would significantly increase chances that Slovenia's aging population will not slip into poverty.

In this paper we also show that allocation of savings is less risky in the context of pensions compared to traditional allocation of savings. Namely, as pension income has to be combined from the PAYG and from the funded pension portfolio, and both depend on specific drivers having different sensitivities to various shocks, pension beneficiaries are protected by diversification benefits. Such benefits are the largest in case of investments in stocks and at times where PAYG is relatively important source of pension income, i.e. when country makes first steps towards building private pension system. Namely, where total pension income already relies to a small extent on the PAYG only pension beneficiaries almost entirely bear the risk of a capital market crash.

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THE PERSISTENCE OF PRICING INEFFICIENCIES IN THE STOCK MARKETS OF THE EASTERN EUROPEAN EU NATIONS

JAMES FOYE[1]
DUŠAN MRAMOR[2]
MARKO PAHOR[3]

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ABSTRACT: *This paper applies a range of metrics to test for the presence of weak form market efficiency in the Eastern European countries that joined the EU in 2004, we test both the years prior to and following accession. The results from our tests indicate that, despite the expectations of many previous studies, even after entering the EU the stock markets of these countries still do not conform to even the loosest form of market efficiency. We improve and extend previous studies by incorporating liquidity controls, applying a wider range of methodologies and by using individual stocks rather than indices.*

Keywords: *Emerging Stock Markets; European Union; Eastern European Transition Countries; Stock Market Efficiency; Weak Form Market Efficiency*
JEL Classification: F36, G14

1. INTRODUCTION

The debate over stock market efficiency is one of the central tenets of capital market theory. The issue is particularly pertinent for the Eastern European nations that joined the European Union in 2004[4] (hereafter the EE EU nations) because of the stock market's role in the ongoing privatization process and also as it serves as an important barometer with which to measure the progress made by these countries in the transition from planned to market economies. In this paper we examine weak form market efficiency (WFME) as defined by Fama (1970) which, as the loosest form of market efficiency, requires nothing more than current period returns "fully reflect" earlier period returns and thus successive price movements are independent of each other: failure to conform to WFME means that stronger forms of efficiency are not present and the stock market's pricing can be considered inefficient.

[1] University of Ljubljana, Faculty of Economics, Slovenia, Ljubljana, e-mail: james.foye@ef.uni-lj.si
[2] University of Ljubljana, Faculty of Economics, Slovenia, Ljubljana, e-mail: dusan.mramor@ef.uni-lj.si
[3] University of Ljubljana, Faculty of Economics, Slovenia, Ljubljana, e-mail: marko.pahor@ef.uni-lj.si

[4] These are the transition nations that joined the EU on 1st May 2004, namely Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia and Slovenia

A significant body of research into WFME in the EE EU nations exists. Jagric et al (2005) test for WFME in the region, the authors found that the stock market indices of Czech Republic, Hungary, Russia and Slovenia all exhibited weak form inefficiencies in the form of long memory in stock returns. Worthington and Higgs (2004) examined WFME in both developed and emerging stock markets in Europe, of the emerging markets covered (Czech Republic, Hungary, Poland and Russia) only the Hungarian stock market could be considered weak form efficient. Gilmoore and McManus (2001) applied a range of WFME tests to the larger EE EU economies (Czech Republic, Hungary and Poland) over the period 1990 to 2000 and found that significant weak form inefficiencies exist in the stock exchanges of all three countries. Chun (2000) reported that while the Hungarian market may be weak-form efficient, the stock markets of the Czech Republic and Poland were inefficient. Nivet (1997) and Gordon and Rittenberg (1995) also found that the Polish stock market could not be considered weak form efficient. Ahmed, Rosser and Uppal (2010) found strong evidence of nonlinear speculative bubbles in Czech Republic, Hungary and Poland. Mihailov and Linowski (2002) and Dezelan (2000) find evidence of weak-form inefficiency in the Latvian and Slovenian stock markets respectively.

We further the above studies in a number of ways. Firstly, we incorporate liquidity controls into our work. It is quite possible that illiquid shares exhibit properties consistent with weak form inefficiency; WFME tests, especially those in emerging markets, need to incorporate liquidity controls in order to ensure that the results are not distorted by apparently predictable returns from infrequently traded securities. In our view this is an omission in the studies listed above that reduces the robustness of results. Indeed, Benic and Franic (2008) found a substantial level of illiquidity in the stock markets of Central and Eastern Europe. Secondly, we include all eight transition countries that acceded to the EU in 2004, while the studies listed above include between one and five of the countries: by considering the region in its entirety, we are able to ascertain a broader and more complete perspective of WFME in the EE EU nations. Thirdly, Jagric, Podobnik and Kolanovics's (2005) dataset ends in 2004, the datasets in the other papers cited end before this. In contrast, our dataset starts in 1999 and runs to the end of 2008. Fourthly, much of the previous work examining WFME in the EE EU nations has been based on stock market indices rather than individual stocks: previously reported findings that the stock markets are inefficient may be due to only a small proportion of the indices' constituents or simply the manner in which the indices are constructed. By using individual stocks, our work provides an important validation of previous work. Furthermore, using individual stocks provides a broader view than using indices alone and may help to provide insight into the underlying causes of the inefficiency. Finally, we use the same metrics as Worthington and Higgs (2004), this is a much broader range than the other cited papers use: our wider range of tests allows us to cross check and validate our results. Furthermore, the results from our work further the existing literature by providing a pre- and post-EU accession comparison.

While the majority of early studies found that returns on the newly-created stock exchanges of the EE EU nations did not conform to WFME, many expected these inefficiencies to disappear over time. Wheeler et al (2002) studied the Warsaw Stock Exchange during its first five years of operation; the authors expected the exchange to become more efficient over time, citing increasing experience of market participants, more sessions per week, more analysts offering better research, and better investor relations departments. Rockinger and Urga (2001) surmised that their finding that the Hungarian market had a lower level of predictability than the markets of Czech Republic, Poland and Russia was partly due to the fact that the Budapest Stock Exchange had operated for a longer period of time. Again, suggesting that the stock markets of the EE EU nations should become more efficient simply due to the passage of time. Moor and Wang (2007) examined the volatility levels on the stock markets of the Czech Republic, Hungary, Poland, Slovenia and Slovakia and concluded that volatility declined as the nations moved into the EU. Worthington and Higgs (2004) hypothesised that there may be a link between the absence of WFME and the small size of some stock markets in the EE EU; this implies pricing efficiency will improve with the growth of these markets. Jagric et al (2005) also proposed a tentative link between a stock market's size and its pricing efficiency. From a macroeconomic perspective, Claessens et al (2000) suggested that EU integration will drive the development process in the EU transition countries. Rapacki and Prochniak (2009) and Vojinovic, Oplotnik and Prochiniak (2010) examined real beta and sigma convergence in the EE EU nations during the process of EU accession, an important extension of this work is to question whether nations' stock markets are also converging as authors such as Csaba (2011, p11) report that "financial institutions play a pre-eminent role in all phases of transformation". Bekaert et al (2013) report that EU membership reduces equity market segmentation.

We test WFME in the EE EU nations over periods 1.1.1999 to 31.12.2003 and 1.1.2004 to 31.12.2008 to determine whether the increasing experience of market participants over time, EU accession and the increasing number of stocks listed, larger market capitalisations and increased turnover in the region has caused markets to become more efficient. Contrary to the expectations of the majority of studies listed above, our tests are all in broad agreement that the equity markets of the EE EU nations do not conform with WFME and this situation has not been substantially affected by accession to the EU. Therefore, none of the factors that previous researchers expected to become catalyst to drive the markets towards higher efficiency have materialized. Despite the passage of almost a generation since the creation of the EE EU stock markets, a significantly larger number of listed securities and 5 years since EU accession, these markets still cannot be considered to conform to WFME: these results pose the question of what changes are needed to improve efficiency of financial markets in these countries or whether these stock exchanges will ever attain pricing efficiency.

2. DATASET

Our dataset consists of stocks included in the Dow Jones Stoxx EU Enlarged Total Market index, using data obtained from Bloomberg. This is a free-float capitalization-weighted index covering the countries have joined the EU since 2004. We excluded stocks from Bulgaria and Romania as the paper is concerned with the countries that joined the EU in

2004. We excluded stocks from Cyprus and Malta as we are only investigating transition countries. Our dataset covers the period from the 1st January 1999 to 31st December 2008, split into subperiods 1st January 1999 to 31st December 3003 (pre-accession) and 1st January 2004 to 31st December 2008 (post-accession). The reason for the use of subperiods lies in the broader range of methodology employed, such as liquidity controls and the use of individual stocks rather than indices, that does not allow direct comparison of our post accession results with previous studies. Although 1st May was the actual accession date, the effects of accession were earlier – this is the reason why we include the entirety of 2004 in our dataset. We did not extend our dataset past 2008 because of the collapse in financial markets.

We use daily Bloomberg last prices and log returns calculated as:

$$\Delta y_{it} = \log(y_{it}) - \log(y_{it-1})$$

Where:
y_{it} = price of stock i at time t

The descriptive statistics for the two datasets are shown in Table 1.

The increasing number of IPOs caused the number of stocks in our post accession dataset to increase to 151 from 97 in our pre accession dataset. As Poland is by far the region's largest economy, it is logical that the country's stock exchange has the largest weight in our dataset; what is interesting is that the number of stocks quoted on the Warsaw Stock Exchange has almost doubled from 55 to 102 between 1999 and 2004, while few new stocks appeared on the other exchanges. Because our dataset contains only a single stock from each of Latvia and Slovakia we are sceptical that we can make any inferences about the stock markets of these countries.

Average returns over the pre-accession period are positive, the financial crisis that began in 2007 resulted in negative returns over the post-accession period. Despite the volatility ensuing from the stock market downturn that began in 2007, the standard deviation of our dataset for 2004-2008 is lower than for 1999-2003, with only Slovenia recording higher volatility. The skewness of our datasets moves from positive to negative, indicating that while over period 1999-2003 there was a greater probability of a large decrease rather than a large increase in stock prices, the opposite was true for period 2004-2008. However, as the skewness readings for 1999-2003 and 2004-2008 are both close to zero, it is hard to draw any firm conclusions. The kurtosis of our dataset decreased significantly between 1999-2003 and 2004-2008, with only the single Latvian stock recording an increase. The Jacque-Bera statistic is used to test the null hypothesis that stock returns are normally distributed. From the associated p-values, it is clear that only stocks listed on the Budapest Stock Exchange over period 1999-2003 could have returns claonsidered to be normally distributed at any conventional level of significance. The results from the Jacque-Bera test are in broad agreement: returns on the stock markets of the EE EU nations are not normally distributed. However, it is clear the Jacque-Bera test is significant

Table 1: *Descriptive Statistic*

	Number of Stocks		Mean		Standard Deviation		Skewness		Kurtosis		Jarque-Bera		Jarque-Bera P Value	
	1999-2003	2004-2008	1999-2003	2004-2008	1999-2003	2004-2008	1999-2003	2004-2008	1999-2003	2004-2008	1999-2003	2004-2008	1999-2003	2004-2008
Entire Region	97	151	0,05%	-0,03%	3,31%	2,64%	-0,01	0,01	34,07	13,27	401.4[illegible]3	14.812	0,00	0,00
Czech Republic	5	7	0,05%	0,00%	2,33%	2,25%	-0,35	-0,61	10,23	15,45	3.443	7.750	0,00	0,00
Estonia	5	8	0,08%	-0,16%	5,05%	2,73%	-4,17	-0,51	141,8[illegible]	10,67	2.620.[illegible]52	4.281	0,00	0,00
Hungary	8	8	0,00%	0,02%	2,43%	2,41%	0,05	-0,02	9,58	8,93	7.[illegible]5[illegible]	2.077	0,04	0,00
Lithuania	10	10	0,17%	0,01%	6,60%	2,42%	1,94	-0,46	89,32	12,81	992.03[illegible]	5.575	0,00	0,00
Latvia	1	1	0,04%	-0,05%	2,77%	2,24%	0,17	0,36	8,43	15,21	1.500	7.636	0,00	0,00
Poland	55	102	0,03%	-0,02%	3,14%	2,79%	0,03	0,10	2[illegible],43	13,21	252.419	16.014	0,00	0,00
Slovakia	1	1	0,08%	0,06%	3,90%	1,81%	-2,38	-0,44	25,08	13,92	14.601	5.511	0,00	0,00
Slovenia	12	14	0,08%	-0,05%	1,60%	2,09%	0,19	0,37	34,74	16,69	161.628	30.662	0,00	0,00

Our dataset is based on stocks in the Dow Jones Stoxx EU Enlarged Total Market index, we only use the transition countries that joined the EU on 1st May 2004.
All calculations are based on daily stock returns calculated on natural logarithms of Bloomberg last prices in local currencies.
Mean is calculated as an arithmetic mean calculated for stocks on an individual basis and then equally weighted for the entire region/individual exchanges
Standard deviation, skewness and kurtosis are calculated on daily returns equally weighted for the entire region/individual exchanges
The skewness and kurtosis inputs for the Jarque-Berra are the same as those reported in this table

due to high kurtosis, rather than skewness, therefore the parametric models we apply still return robust results.

3. METHODOLOGY

The tests we employ fall into four categories: tests of serial independence, unit root tests, multiple variance ratio tests and liquidity. We chose to replicate the methodology of Worthington and Higgs (2004) for the serial independence, unit root tests and multiple variance ratio tests because of the broad range of WFME tests applied by the authors and the recognition it received in the literature. Griffin et al (2010) use a range of tests similar to the ones employed in this paper, they question whether it is feasible to make statements about relative market efficiency internationally unless one can control for the information environment. While our dataset covers nations with substantial similarities, we cannot rule out the possibility that our results have been distorted by this. While our dataset covers a large geographic area, the majority of stocks are quoted on the Warsaw Stock Exchange. To control for any Polish bias, we perform the tests for both the region as a whole and the individual countries.

3. 1 Tests of serial independence

A time series is said to be serially correlated if a regression of a time series of returns with its own lags yields statistically significant results:

$$E(\Delta y_{it}|\Delta y_{it-1}) = \beta_1 + \beta_2 \Delta y_{it-1}$$

Where:

$E(\Delta y_{it}|\Delta y_{it-1})$ = the expected value of Δy_{it} given Δy_{it-1}

β_1 = the regression intercept

β_2 = the regression slope

Unlike serial correlation, the runs test is non-parametric and therefore does not require the returns to be normally distributed. Runs tests determine whether a time series follows a random walk by counting the number of consecutive positive or negative observations and comparing it to an expected value *(E(R))*:

$$E(R) = \frac{N + 2N_U N_D}{N}$$

Where:
N = Number of observations

N_U = *Number of positive observations*
N_D = *Number of negative observations*
R = *Number of runs*

We use the expected value and variance values (V(R)) to calculate a test statistic, Z:

$$V(R) = \frac{2N_u N_D(2N_D N_D - N)}{(N)^2(N-1)}$$

$$Z = \frac{R - E(R)}{\sqrt{V(R)}}$$

The null hypothesis is that the returns can be considered to follow a random walk process. Rejection of the null hypothesis indicates that the stock's returns are non-random and contravene WFME. In order to test whether EU accession resulted in an increase in WFME, we use a z-test to determine if the percentage of stocks considered statistically significant at a particular significance level is statistically different between the pre- and post-accession datasets.

3.2 Unit root tests

Unit root tests are used to determine whether the log returns of stocks in our dataset is stationary, i.e. whether it has constant statistical properties; if stocks follow a random walk process, stock returns should be non-stationary. We use three variants, Augmented Dickey Fuller (ADF), Phillips-Perron (PP) and Kwaitkowki, Phillips, Schmidt and Shin (KPSS).

ADF is the most well-known unit root test, the null hypothesis is that the data is nonstationary. The measure is calculated by running the following regression:

$$\Delta y_{it} = \beta_0 + \beta_1 tr + \alpha_0 yi_{t-1} + \alpha_p \sum_{p=1}^{q} \Delta y_{it-p} + \varepsilon_{pt}$$

Where:
α = the coefficients to be estimated
q = number of lagged terms
β_0 = intercept
β_1 = trend coefficient
tr = trend

MacKinnon's critical values are then applied to determine the significance of α.

The PP test, developed by Phillips and Perron (1988), extends ADF to allow errors to be independent and heteroscedastic. For a complete derivation, see Phillips and Perron (1988).

While the ADF and PP tests have null hypothesis of nonstationarity, the KPSS test has a null hypothesis of stationarity. Reversing the null hypothesis provides a useful validation check for the results from the ADF and PP tests. The reader should consult Kwiatkowski et al. (1992) for a full derivation. As with the tests of serial independence, we apply a z-test to determine whether the results from the pre- and post-accession datasets can be considered statistically different.

3.3 Multiple Variance Ratio Tests

The third set of statistics employed are multiple variance ratio (MVR) tests. This approach was developed by Lo and MacKinlay (1988, 1989) and Chow and Denning (1993) who constructed the MVR tests in order to detect both autocorrelation and heteroscedasticity in returns. This is important because if stocks follow a random walk, the variance of returns should rise as a linear function to the number of observations. That is, the variance ratio of the returns over qq period must be equal to $q\sigma^2 q\sigma^2$. The variance ratio (VR) is calculated as:

$$VR(q) = \frac{\sigma^2(q)}{\sigma^2(1)}$$

Where:
$\sigma^2(1)$ = variance of daily log returns
q = number of periods used for the sampling interval
$\sigma^2(q) = (1/q)$ multiplied by the variance of q-daily returns

If stocks conform to the random walk process, VR should not be statistically different to one. In line with the methodology of Worthington and Higgs (2004), the sampling intervals used for q were 2, 5, 10 and 20 days. For a more in depth overview of MVR methodology or a complete derivation, the reader should consult Worthington and Higgs (2004) or Chow and Denning (1993) respectively. We also apply a z-test to determine whether the pre- and post-accession results are statistically different.

3.4 Liquidity Controls

Studies frequently conclude that liquidity is related to future returns. Examples of such work include Amihud and Mendelson (1986, 1989), Chordia et al (2001), Jones (2002), Amihud (2002), and Brennan et al (1998). Datar et al (1998) demonstrate a negative correlation between liquidity, as measured by turnover, and returns. Haugen and Baker (1996) found that liquidity is one of several generic factors that explain returns across global stock markets. Brzeszczynski et al (2011) found that trading intensity affected beta calculations for stocks listed on the Warsaw Stock Exchange and thus had serious ramifications for corporate finance decisions.

The relatively small size of the stock markets of the EE EU countries raises the concern that our results could be distorted by liquidity issues. Liquidity is an elusive concept, consequently in Table 5 we employ three widely used measures to control for it: i) Market capitalization ii) Average volume divided by shares outstanding iii) Bid-ask spread divided by share price. We create liquidity portfolios by assigning a rank (1 (low) to 5 (high)) to every stock for each of the three liquidity measures. Then we separate the combined results from Tables 2, 3, and 4 into five liquidity ranked portfolios in order to examine the effects of liquidity on the tests employed; we repeat this for each of market capitalization (Panel A) average volume divided by shares outstanding (Panel B) and Bid-ask spread (Panel C).

4. RESULTS

The results from the tests of serial independence, unit root tests and multiple variance ratio tests are shown in Tables 2, 3 and 4 respectively. As we cover a large geographic region, each table also provides a geographic breakdown of the results. While around one-third of our dataset is listed outside Poland, the shares are listed on a lot of different exchanges; no exchange other that the Warsaw Stock Exchange has more than 14 shares in the dataset. This makes inferences for individual countries difficult.

4.1 Tests of serial independence

Table 2 shows the results from the tests of serial independence, the serial correlation coefficient and the runs test.

Looking at all the stock exchanges in the dataset, even at the 0.01 level of significance, almost one third of the stocks in our dataset return significant t-statistics from the serial correlation regressions for both the pre- and post-EU accession periods. Whilst there has been a marginal decrease in the number of stocks statistically significant at the 0.01 level between the pre- and post-accession datasets, the z-test reveals that the difference is not statistically significant. 43% of stocks in our dataset can be considered serially correlated at the 0.1 significance level for the pre-accession period; this rises to 66% for the post-accession period. The z-test reveals that the increase in the number of stocks exhibiting serial correlation at the 0.05 and 0.1 levels is statistically significant at 0.01, indicating that prices of stocks listed in the EE EU nations may have actually become less efficient. Looking at the individual stock exchanges, it can be seen that the results from the stock exchanges of other countries are largely consistent with those from the Warsaw Stock Exchange. Across the majority of stock exchanges most stocks exhibit properties consistent with serial correlation, at least at the 0.1 level. The z-test reveals no statistically significant difference between the pre- and post-accession datasets. Thus we can comfortably reject the null hypothesis that returns in the stock markets of the EE EU are not serially correlated.

Table 2: *Tests of Serial Independence*

		Serial Correlation T Statistic			Runs Test		
		1999-2003	2004-2008	Z Test	1999-2003	2004-2008	Z Test
Entire Region							
% of Observations Significant at	1%	31%	28%	0,41	22%	19%	0,47
	5%	39%	54%	- 2,23	38%	38%	- 0,04
	10%	43%	66%	- 3,45	46%	49%	- 0,40
% of Negative Observations		15%	42%		64%	69%	
Czech Republic							
% of Observations Significant at	1%	40%	29%		20%	0%	
	5%	60%	43%		40%	29%	
	10%	60%	71%		60%	43%	
% of Negative Observations		0%	29%		80%	57%	
Estonia							
% of Observations Significant at	1%	60%	50%		40%	38%	
	5%	60%	63%		60%	50%	
	10%	60%	63%		60%	63%	
% of Negative Observations		60%	75%		20%	63%	
Hungary							
% of Observations Significant at	1%	13%	38%		13%	38%	
	5%	38%	50%		13%	63%	
	10%	50%	63%		13%	75%	
% of Negative Observations		38%	50%		13%	38%	
Latvia							
% of Observations Significant at	1%	0%	100%		0%	0%	
	5%	0%	100%		0%	0%	
	10%	0%	100%		0%	0%	
% of Negative Observations		0%	0%		100%	0%	
Lithuania							
% of Observations Significant at	1%	40%	70%		70%	30%	
	5%	70%	90%		90%	50%	
	10%	70%	90%		90%	70%	
% of Negative Observations		30%	20%		90%	90%	
Poland							
% of Observations Significant at	1%	13%	16%		11%	18%	
	5%	16%	48%		27%	35%	
	10%	22%	63%		40%	44%	
% of Negative Observations		11%	48%		71%	74%	
Slovakia							
% of Observations Significant at	1%	100%	0%		100%	100%	
	5%	100%	0%		100%	100%	
	10%	100%	0%		100%	100%	
% of Negative Observations		0%	0%		0%	100%	
Slovenia							
% of Observations Significant at	1%	100%	71%		25%	7%	
	5%	100%	71%		50%	36%	
	10%	100%	71%		50%	50%	
% of Negative Observations		0%	0%		58%	50%	

All calculations are based on stock returns calculated on natural logarithms of Bloomberg last prices in local currencies.
Serial correlation is calculated using one day lags
Runs tests calculations are based on the sign of returns

When the runs test was applied to our dataset, about one fifth of stocks yielded statistically significant results even at the most stringent 0.01 level for both the 1999-2004 and 2004-2008 datasets. Around half of both the pre- and post-accession datasets can be considered significant at the 0.1 level. Stocks listed on the Riga Stock Exchange perform poorly in the runs tests, but the dataset only contains one stock from this country; excluding Latvia, the non-Polish stock markets have similar results to the entire dataset.

4.2 Unit root tests

Table 3 shows the results from the three sets of statistics that form the unit root tests. The null hypothesis of the ADF and PP tests is that the time series has a unit root. The KPSS test reverses the null hypothesis and assumes that the time series has no unit root.

Both the ADF and PP tests reject the null hypothesis, even at 0.01, for all stocks in both the pre- and post-accession datasets. We can comfortably reject the null hypothesis of nonstationarity for all stocks. Needless to say, there is no country variation here. Both tests clearly indicate that the returns of all stocks in the dataset are stationary, that is follow a deterministic rather than stochastic trend; inconsistent with a random walk.

Out of all the metrics we employ, only the KPSS test indicates that stationarity may have declined between the pre- and post-accession periods. The KPSS statistic is insignificant for less than half of all stocks at the 0.01 level of significance for the post-accession dataset, indicating that we cannot reject the null hypothesis of no unit root; yet for our pre-accession dataset, only 5% of stocks have KPSS statistics that can be considered statistically significant at the 0.01 level. Whilst almost three quarters of post-accession stocks have KPSS statistics that can be considered statistically significant at 0.1, the corresponding figure for the pre-accession nations is only around one quarter. The z-test reveals that there is a statistically significant increase in the KPSS statistic between the pre- and post-accession datasets. The results from Poland are almost identical to those for the region as a whole, indicating little regional variation.

While the KPSS statistic is less conclusive than ADF or PP, we can still confidently infer that all three unit root tests employed indicate that returns of many stocks listed in the EE EU nations are stationary, leading us to reject the null hypothesis that stocks follow a random walk.

4.3 Multiple Variance Ratio Tests

Table 4 shows the results from the MVR tests using sampling intervals of two days, 5 five days, 10 days and 20 days; corresponding to one day, one week, one fortnight and one month.

Table 3: *Unit Root Tests*

		ADF		Phillips-Perron Test		KPSS Test		
		1999-2003	2004-2008	1999-2003	2004-2008	1999-2003	2004-2008	Z Test
Entire Region								
% of Observations	1%	100%	100%	100%	100%	5%	46%	- 6,81
Significant at	5%	100%	100%	100%	100%	13%	64%	- 7,86
	10%	100%	100%	100%	100%	25%	72%	- 7,31
	Average	-29,27	-28,88	-33,76	-31,09	0,26	0,79	
	Absolute Average	29,27	28,88	33,76	31,09	0,26	0,79	
Czech Republic								
% of Observations	1%	100%	100%	100%	100%	0%	0%	
Significant at	5%	100%	100%	100%	100%	0%	14%	
	10%	100%	100%	100%	100%	20%	29%	
	Average	-31,61	-28,49	-33,00	-32,15	0,16	0,34	
	Absolute Average	31,61	28,49	33,00	32,15	0,16	0,34	
Estonia								
% of Observations	1%	100%	100%	100%	100%	0%	63%	
Significant at	5%	100%	100%	100%	100%	0%	75%	
	10%	100%	100%	100%	100%	20%	75%	
	Average	-31,57	-28,50	-34,43	-30,91	0,24	0,88	
	Absolute Average	31,57	28,50	34,43	30,91	0,24	0,88	
Hungary								
% of Observations	1%	100%	100%	100%	100%	0%	50%	
Significant at	5%	100%	100%	100%	100%	0%	75%	
	10%	100%	100%	100%	100%	0%	75%	
	Average	-31,32	-31,86	-31,37	-35,00	0,13	0,65	
	Absolute Average	31,32	31,86	31,37	35,00	0,13	0,65	
Latvia								
% of Observations	1%	100%	100%	100%	100%	0%	100%	
Significant at	5%	100%	100%	100%	100%	0%	100%	
	10%	100%	100%	100%	100%	0%	100%	
	Average	-36,60	-35,06	-36,58	-35,10	0,27	1,04	
	Absolute Average	36,60	35,06	36,58	35,10	0,27	1,04	
Lithuania								
% of Observations	1%	100%	100%	100%	100%	30%	90%	
Significant at	5%	100%	100%	100%	100%	60%	100%	
	10%	100%	100%	100%	100%	60%	100%	
	Average	-18,52	-25,24	-30,20	-32,16	0,56	1,73	
	Absolute Average	18,52	25,24	30,20	32,16	0,56	1,73	
Poland								
% of Observations	1%	100%	100%	100%	100%	4%	39%	
Significant at	5%	100%	100%	100%	100%	9%	60%	
	10%	100%	100%	100%	100%	22%	69%	
	Average	-30,73	-28,75	-33,88	-30,30	0,23	0,71	
	Absolute Average	30,73	28,75	33,88	30,30	0,23	0,71	
Slovakia								
% of Observations	1%	100%	100%	100%	100%	0%	100%	
Significant at	5%	100%	100%	100%	100%	0%	100%	
	10%	100%	100%	100%	100%	0%	100%	
	Average	-22,35	-34,85	-22,24	-34,81	0,11	0,96	
	Absolute Average	22,35	34,85	22,24	34,81	0,11	0,96	
Slovenia								
% of Observations	1%	100%	100%	100%	100%	0%	64%	
Significant at	5%	100%	100%	100%	100%	17%	79%	
	10%	100%	100%	100%	100%	33%	93%	
	Average	-28,19	-30,21	-38,55	-32,87	0,27	0,93	
	Absolute Average	28,19	30,21	38,55	32,87	0,27	0,93	

All calculations were made on natural logarithms of Bloomberg last prices in local currency
Augmented Dickey Fuller (ADF) test, H0: unit root, H1: no unit root (stationary)
Phillips Peron (PP), H0: unit root, H1: no unit root (stationary)
Kwiatkowski, Phillips, Schmidt and Shin (KPSS), H0: no unit root (stationary), H1: unit root

Table 4: *Multiple variance ratio tests*

		T Statistic q=2			T Statistic q=5			T Statistic q=10			T Statistic q=20			Stocks significant at at least one of the above time intervals		
		1999-2003	2004-2008	Z Test	1999-2003	2004-2008	Z Test	1999-2003	2004-2008	Z Test	1999-2003	2004-2008	Z Test	1999-2003	2004-2008	Z Test
Entire Region																
% of Observations	1%	14%	19%	-0,97	12%	21%	-1,66	7%	19%	-2,62	9%	20%	-2,24	22%	31%	-1,63
	5%	31%	32%	-0,25	26%	32%	-1,12	20%	28%	-1,47	16%	29%	-2,27	41%	44%	-0,38
Significant at	10%	38%	38%	-0,04	35%	38%	-0,43	25%	33%	-1,41	25%	37%	-2,03	55%	54%	0,05
Czech Republic																
% of Observations	1%	0%	14%		0%	0%		0%	0%		0%	0%		0%	14%	
	5%	40%	14%		20%	14%		20%	0%		0%	0%		40%	14%	
Significant at	10%	40%	14%		20%	14%		20%	0%		0%	0%		40%	14%	
Estonia																
% of Observations	1%	40%	25%		40%	0%		20%	0%		20%	13%		60%	38%	
	5%	60%	50%		80%	38%		80%	0%		60%	13%		100%	75%	
Significant at	10%	80%	50%		80%	38%		80%	13%		80%	25%		100%	88%	
Hungary																
% of Observations	1%	13%	13%		13%	13%		0%	0%		13%	0%		25%	13%	
	5%	13%	25%		13%	13%		13%	13%		13%	0%		25%	25%	
Significant at	10%	13%	38%		25%	13%		25%	13%		50%	13%		63%	38%	
Latvia																
% of Observations	1%	0%	0%		0%	0%		0%	0%		0%	0%		0%	0%	
	5%	0%	0%		0%	0%		0%	0%		0%	0%		0%	0%	
Significant at	10%	0%	0%		0%	0%		0%	0%		0%	0%		0%	0%	
Lithuania																
% of Observations	1%	10%	10%		20%	30%		20%	30%		30%	60%		40%	60%	
	5%	30%	40%		30%	40%		50%	50%		40%	70%		60%	70%	
Significant at	10%	30%	40%		50%	40%		50%	50%		40%	80%		70%	80%	
Poland																
% of Observations	1%	15%	22%		11%	25%		5%	25%	-2,98	5%	22%		18%	31%	
	5%	27%	34%		24%	37%		13%	33%	-2,80	13%	33%		35%	45%	
Significant at	10%	38%	41%		35%	44%		16%	39%	-2,95	18%	41%		49%	56%	
Slovakia																
% of Observations	1%	0%	0%		0%	0%		0%	0%		0%	0%		0%	0%	
	5%	0%	0%		0%	0%		0%	0%		0%	0%		0%	0%	
Significant at	10%	0%	0%		0%	0%		0%	0%		0%	0%		0%	0%	
Slovenia																
% of Observations	1%	17%	14%		8%	14%		8%	7%		8%	7%		17%	29%	
	5%	50%	21%		25%	14%		8%	14%		8%	14%		50%	29%	
Significant at	10%	50%	29%		25%	21%		25%	21%		17%	21%		58%	43%	

All calculations were made on natural logarithms of Bloomberg last prices in local currency; Sampling intervals (q) are in days

Even at the 0.01 level of significance, the MVR tests generally suggest that many stocks in our dataset do not follow a random walk process. While the percentage of stocks significant for at least one of the q levels is substantially higher for the post-accession dataset than the pre-accession dataset, the z-tests reveal that this is not statistically significant. At the 0.1 level of significance, more than half of all stocks do not conform to a random walk process for at least one of the sampling intervals applied, and the results are very similar for the pre- and post-accession nations. Excluding Czech Republic, Latvia and Slovakia (the small number of stocks listed in these nations makes inferences about them questionable anyway), there is not a large variation amongst the different countries in our dataset, with the results for Poland and the entire region being almost identical.

4.4 Liquidity Controls

Table 5 shows the results from the liquidity controls employed:

The results from using market capitalization as a proxy for liquidity are shown in Table 5 Panel A. For both the pre- and post-accession datasets, smaller capitalized stocks exhibit higher levels of serial correlation. Runs tests are also substantially affected by their market capitalization quintile, with the smaller market capitalization quintile stocks returning a higher proportion of significant results. The ADF and PP tests are both excluded from the table as every stock in our dataset can be considered statistically significant at the 0.01 level and thus there is no variation across any of the liquidity quintiles. For the KPSS tests, the results for the large market capitalization quintile are very similar to those from the small market capitalization quintile, therefore there is nothing to suggest that the KPSS tests is affected by liquidity (as measured by market capitalization). For the MVR tests, portfolio 5 actually has a higher percentage of stocks returning statistically significant results than any of the other four quintiles: lack of liquidity is clearly not distorting results from the MVR tests. Whilst lack of liquidity associated with smaller market capitalization may have distorted some of the tests of serial independence, a substantial number of stocks in the largest market capitalization portfolio still return significant results. Market capitalization does not have any meaningful effect on any of the three unit root tests of the MVR tests.

The results from using average volume divided by shares outstanding as a liquidity control are shown in Table 5 Panel B. For serial correlation, the number of stocks significant at each of the three significance levels we use is actually higher in the most liquid portfolio 5 than in the least liquid portfolio 1. Therefore, there is no indication that lack of liquidity, as measured by average volume divided by shares outstanding, is distorting the serial correlation tests. Whilst the runs tests return the highest percentage of significant results for the lowest-liquidity portfolio 1, there is not a huge amount of variation across the quintiles. In a similar manner to the serial correlation statistic, the percentage of stocks returning significant results for the KPSS tests actually increases as liquidity increases. The MVR tests return very similar results across the five quintiles. It is clear that

Table 5 Panel A: *Liquidity Controls - Market Cap*

		Tests of Serial Independence						Unit Root Tests		Multiple variance ratio tests	
		Serial Correlation T Statistic		Number of Runs		Runs Test			KPSS Test	Stocks significant for at least one of the q sampling intervals	
		2000-2004	2004-2008	2000-2004	2004-2008	2000-2004	2004-2008	2000-2004	2004-2008	2000-2004	2004-2008
Quintile 1											
Observations %	1%	50%	11%			30%	26%	10%	51%	30%	34%
of Observations	5%	60%	86%			45%	46%	20%	69%	55%	60%
Significant at	10%	65%	89%			50%	63%	20%	80%	75%	69%
	Average			422,70	515,03						
Quintile 2											
Observations %	1%	33%	15%			19%	15%	3%	35%	19%	27%
of Observations	5%	33%	31%			39%	31%	11%	50%	36%	42%
Significant at	10%	42%	77%			42%	38%	19%	54%	53%	54%
	Average			540,03	496,62						
Quintile 3											
Observations %	1%	50%	73%			0%	27%	0%	53%	50%	37%
of Observations	5%	50%	80%			50%	50%	0%	63%	50%	43%
Significant at	10%	50%	90%			50%	53%	0%	70%	50%	57%
	Average			319,00	512,00						
Quintile 4											
Observations %	1%	23%	33%			9%	10%	5%	33%	14%	17%
of Observations	5%	41%	50%			27%	27%	5%	60%	32%	20%
Significant at	10%	41%	50%			45%	50%	23%	77%	36%	37%
	Average			478,77	438,63						
Quintile 5											
Observations %	1%	12%	10%			35%	17%	6%	53%	24%	40%
of Observations	5%	24%	13%			41%	37%	24%	77%	47%	50%
Significant at	10%	24%	20%			53%	37%	47%	77%	59%	53%
	Average			524 47	552,57						

Market capitalization is an average of daily market capitalization taken from Bloomberg in Euros over the period for which the weak form tests for the individual security were calculated.
Stocks are ranked according to their market capitalization in Euros and assigned to quintiles 1 (low) to 5 (high). The results from Tables 2, 3 and 4 are shown for each market capitalization quintile.

Table 5 Panel B: *Liquidity Controls – Average Volume Divided by Shares Outstanding*

		Tests of Serial Independence						Unit Root Tests		Multiple variance ratio tests	
		Serial Correlation T Statistic		Number of Runs		Runs Test		KPSS Test		Stocks significant for at least one of the q sampling intervals	
		2000-2004	2004-2008	2000-2004	2004-2008	2000-2004	2004-2008	2000-2004	2004-2008	2000-2004	2004-2008
Quintile 1											
Observations %	1%	35%	16%			45%	26%	15%	48%	45%	35%
of Observations	5%	50%	29%			65%	42%	30%	55%	65%	52%
Significant at	10%	50%	71%			70%	58%	40%	55%	80%	65%
	Average			395,35	527,39						
Quintile 2											
Observations %	1%	5%	0%			11%	20%	5%	53%	16%	37%
of Observations	5%	16%	7%			26%	30%	16%	73%	32%	53%
Significant at	10%	16%	10%			47%	40%	47%	73%	37%	53%
	Average			530,21	477,07						
Quintile 3											
Observations %	1%	84%	37%			21%	7%	0%	37%	11%	20%
of Observations	5%	95%	57%			37%	30%	11%	60%	37%	23%
Significant at	10%	95%	60%			47%	47%	21%	80%	42%	40%
	Average			527,89	497,43						
Quintile 4											
Observations %	1%	32%	7%			11%	23%	0%	43%	11%	27%
of Observations	5%	32%	87%			32%	43%	5%	67%	47%	53%
Significant at	10%	42%	87%			37%	53%	5%	80%	58%	63%
	Average			595,89	513,03						
Quintile 5											
Observations %	1%	0%	83%			20%	20%	5%	47%	25%	37%
of Observations	5%	5%	90%			30%	47%	5%	67%	25%	37%
Significant at	10%	15%	100%			30%	47%	10%	73%	55%	50%
	Average			432,45	501,97						

Both average volume and shares outstanding were calculated as average of daily observations taken from Bloomberg over the period for which the weak form tests for the individual security were calculated.
Stocks are ranked according to average volume divided by shares outstanding and assigned to quintiles 1 (low) to 5 (high). The results from Tables 2, 3 and 4 are shown for each average volume divided by shares outstanding quintile.

Table 5 Panel C: *Liquidity Controls – Bid-Ask Spread*

		Tests of Serial Independence						Unit Root Tests		Multiple variance ratio tests	
		Serial Correlation T Statistic		Number of Runs		Runs Test		KPSS Test		Stocks significant for at least one of the q sampling intervals	
		2000-2004	2004-2008	2000-2004	2004-2008	2000-2004	2004-2008	2000-2004	2004-2008	2000-2004	2004-2008
Quintile 1											
Observations %	1%	0%	58%			13%	10%	0%	45%	21%	26%
of Observations	5%	4%	74%			33%	26%	4%	65%	33%	29%
Significant at	10%	17%	74%			38%	35%	8%	77%	50%	42%
	Average			530,96	521,19						
Quintile 2											
Observations %	1%	60%	60%			27%	20%	7%	43%	27%	30%
of Observations	5%	60%	73%			40%	47%	7%	67%	47%	37%
Significant at	10%	60%	80%			40%	53%	13%	80%	53%	57%
	Average			358,67	452,07						
Quintile 3											
Observations %	1%	84%	0%			32%	20%	5%	42%	26%	28%
of Observations	5%	89%	50%			47%	40%	21%	54%	53%	48%
Significant at	10%	89%	76%			58%	56%	26%	62%	63%	58%
	Average			542,16	502,82						
Quintile 4											
Observations %	1%	22%	50%			22%	30%	9%	50%	17%	30%
of Observations	5%	39%	70%			39%	60%	17%	60%	39%	60%
Significant at	10%	43%	80%			48%	80%	30%	60%	52%	70%
	Average			501,30	543,50						
Quintile 5											
Observations %	1%	0%	7%			19%	23%	6%	53%	19%	43%
of Observations	5%	13%	13%			31%	33%	19%	80%	38%	53%
Significant at	10%	13%	20%			50%	37%	50%	80%	56%	53%
	Average			501,75	524,63						

Bid-ask spread and last price is an average of daily market capitalization taken from Bloomberg in Euros over the period for which the weak form tests for the individual security were calculated. The bid-ask spread was divided by the last price in order to obtain a percentage measure.
Stocks are ranked according to their Bid-ask spread and assigned to quintiles 1 (low) to 5 (high). The results from Tables 2, 3 and 4 are shown for each Bid-ask spread quintile.

liquidity as measured by average volume divided by shares outstanding is not distorting any of the results from these tests.

The results from using the bid-ask spread as a liquidity measure are shown in Table 5 Panel C. Note that, unlike the market capitalization and average volume divided by shares outstanding liquidity controls, higher bid-ask spreads are associated with lack of liquidity. Thus portfolio 1 contains stocks with the lowest bid-ask spreads and highest liquidity. For the serial correlation tests, quintile 1 returns a greater percentage of stocks with statistically significant results than quintile 5; therefore, lack of liquidity is not distorting these results. For the runs tests, the extreme bid-ask portfolios 1 and 5 return the lowest percentage of statistically significant results for the runs test; the median quintile 3 returns the highest percentage of statistically significant results: runs test results are not affected by liquidity as measured by bid-ask spread. The KPSS tests return a marginally higher percentage of statistically significant results for quintile 5, but the results are largely consistent across quintiles. The numbers of stocks returning statistically significant results from the MVR tests increases for the wider bid-ask quintiles, but the lower bid-ask quintiles still return a substantial number of statistically significant results. We can thus conclude that bid-ask spread is not distorting the results of our WFME tests.

Hence from the liquidity tests employed it is clear that the apparent weak-form inefficiencies highlighted by the WFME tests cannot be entirely explained away by liquidity issues. While liquidity may have some explanatory power for some of the tests, it is clear that lack of liquidity is not creating a spurious sense of weak form inefficiency.

5. CONCLUDING REMARKS

The tests employed are in broad agreement: the stock markets of the EE EU nations are not WFME, nor have they become more efficient since EU accession. This contravenes the expectations of many academics who expected these markets to become more efficient and leads us to hypothesize that the inefficiencies will remain for years to come. Many researchers suggested that the passage of time would allow market participants to gain experience and make markets more efficient, however as this has not happened after nearly 20 years of operating, there is no reason to presume that it ever will. Some earlier studies argued that the process of EU integration will improve market efficiency, however our dataset covers the 5 years following EU accession and these markets are still inefficient. Finally, some suggested that the small size of the stock markets of the EE EU made them inefficient, however the number of stocks listed on the Warsaw Stock Exchange has increased to make the number of listed shares comparable to the exchanges of the pre-enlargement EU nations, yet our results show that the Polish stock market is no more efficient than the rest of the EE EU region. The reasons researchers gave for expecting the stock markets of the EE EU nations to become WFME have clearly not materialized: given this, it is hard to see what catalyst can drive these markets to become efficient. Therefore we expect the stock markets of the EE EU countries to remain weak form inefficient for the foreseeable future.

Given our tests incorporate two sub periods and indicate no improvement in the level of WFME in the EE EU nations, our view is that these stock markets will take a significant amount of time to show any meaningful improvement in WFME. This has substantial ramifications. While the issue is most obviously of interest to researchers and market participants engaged in technical analysis and trading models, lack of WFME also has much more important implications for corporate financial decisions and the development of the broader economy. There is a well-established link between pricing efficiency and the efficient allocation of capital; consequently, the absence of WFME in the EE EU nations may impair corporate finance decisions and prevent companies from attaining an optimal capital structure. Even more importantly, the link between the pricing efficiency of a country's stock market and the nation's overall economic development and the possibility that the availability of stock market financing can enhance economic growth means that it is clear that WFME has significant ramifications not just for a country's capital market but also its overall economic development. Furthermore, WFME is of particular importance in the EE EU countries: an efficient capital market can facilitate the ongoing privatization process; as these nations are aiming for economic convergence with the pre-enlargement EU nations, the stock market clearly has a large role to play here; finally, as Worthington and Higgs (2004) suggest, the absence or presence of WFME in Europe's developing markets is an important consideration in the debate about what technological and regulatory reform is necessary or even whether the region's exchanges should merge.

In this paper we focussed on establishing whether stocks listed in the EE EU conform with WFME. Although we do not offer a concrete explanation of the causes of WFME in the region's stock markets, previous research such as Griffin et al (2010) focused on transaction costs and information flow as the drivers of WFME. Thus policy makers may consider improving information flow for example through access to real time prices or encouraging research coverage from a wide range of analysts, alternatively lowing transaction costs may offer a solution.

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EARLY WARNING MODELS FOR SYSTEMIC BANKING CRISES IN MONTENEGRO

ŽELJKA ASANOVIĆ[1]

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ABSTRACT: *The purpose of this research is to create an adequate early warning model for systemic banking crises in Montenegro. The probability of banking crisis occurrence is calculated using discrete dependent variable models, more precisely, estimating logit regression. Afterwards, seven simple logit regressions that individually have two explanatory variables are estimated. Adequate weights have been assigned to all seven regressions using the technique of Bayesian model averaging. The advantage of this technique is that it takes into account the model uncertainty by considering various combinations of models in order to minimize the author's subjective judgment when determining reliable early warning indicators. The results of Bayesian model averaging largely coincide with the results of a previously estimated dynamic logit model. Indicators of credit expansion, thanks to their performances, have a dominant role in early warning models for systemic banking crises in Montenegro. The results have also shown that the Montenegrin banking system is significantly exposed to trends on the global level.*

Key words: *early warning systems, systemic banking crises, logit model, Bayesian model averaging, credit expansion, Montenegro*

JEL Classification: G01; C25; C11

1 INTRODUCTION

Considering high costs of resolving systemic banking crises and their significant negative effects on the economy and therefore on the standard of living, it is necessary to dedicate a lot of attention to research on how and why crises happen in order to try to predict them. Neither are the most developed economies spared of financial crises, including banking, currency and debt crises. The global economic crisis that has started as the US mortgage market crisis unequivocally shows that even developed economies do not pay enough attention to early warning models for systemic banking crises. Namely, these models, even when implemented, are not adequately used. Also, nowadays when there are very significant interdependencies between financial markets, consequences might hardly stay only within the borders of countries hit by the financial crisis. Unlike nowadays, during previous decades these models related mostly to currency crises since currency crises used to occur more often than banking crises.

[1] University of Montenegro, Faculty of Economics; Central Bank of Montenegro, e-mail: zeljka.asanovic@cbcg.me

Extensive empirical literature indicates that, in general, there are two approaches for designing early warning systems that are most commonly used. The first one is a signal approach (non-parametric) that studies and compares behavior of economic indicators for the period before and during the crisis. This approach developed by Kaminsky & Reinhart (1996), and Kaminsky, Lizondo & Reinhart (1998), is also known as the KLR method. The second approach (parametric) calculates the probability of banking crisis occurrence using discrete dependent variable models, estimating usually probit or logit regression (Demirgüç-Kunt & Detragiache, 1998; Eichengreen & Rose, 1998). Besides logit regression, the Bayesian model averaging technique is also applied in this paper. Bayesian model averaging (BMA) takes into account the model uncertainty by taking into consideration various combinations of models, and therefore it enables the author's subjective judgment to be minimized when determining reliable early warning indicators.

The basic motive for this research is a great importance that early warning models have, primarily for the stability of the banking system, as well as for the entire financial system of a country. There is no early warning model for banking crises in Montenegro. One of the main characteristics of the Montenegrin financial system is its relatively simple structure that is a common feature of many developing countries. Banks have a dominant role in the financial system; primarily in financing the private sector because it doesn't have enough own funds accumulated. The banking sector that consists of eleven banks and six microcredit financial institutions is based on the traditional banking. Development of the Montenegrin banking sector during the pre-crisis period is characterized with enormously high credit growth rates. Montenegro was one of European developing countries with the fastest economic growth. Therefore, in 2007 its economic growth reached the peak of 10.70%, while the lowest growth rate was -5.70% in 2009.

Economic slowdown and sudden stop of credit activity supported by the global economic crisis has led to much more deepening of the crisis in Montenegro. A significant problem is that some borrowers are not able to repay regularly loans approved mostly during the credit expansion. One of the reasons due to which the Montenegrin economy has found itself in a very unfavorable situation is the reduced intermediation function of Montenegrin banks. The banking crisis and later also the economic crisis have caused the deterioration of the fiscal position of Montenegro. One of the most significant consequences of the crisis is an intensive growth of sovereign debt. Namely, during 2010 and 2011, the emission of Euro bonds in the amount of EUR 380 million contributed largely to the growth of sovereign debt. In order to prevent a scenario like this to happen again, it is necessary to create and implement early warning models for systemic banking crises.

2 METHODOLOGY AND AVAILABILITY OF DATA

When compared with the signal approach, the advantage of the logit model is that it enables estimation of all variables simultaneously. However, unlike the signal approach, using this method it is not possible to rank indicators according to their relative prognostic power in predicting systemic banking crises. Ranking indicators according to their deviation from the normal behavior would be of a great help to monetary policy holders, because they could determine more easily what corrective measures would be necessary.

This shortcoming might be partially overcome using Bayesian model averaging. Namely, applying this technique it is possible to assign adequate weights to simple logit models with at most two explanatory variables. Although individual variables do not have weights, their relative importance can be approximately determined on the basis of weights assigned to the model that contains these variables.

Logit regression is used in this paper in the same manner as in the most papers dealing with early warning systems for banking crises. The observed time period is divided into two periods: the signal horizon where the dependent variable takes the value 1, and the period out of the signal horizon where the dependent variable takes the value 0. However, there is a difference between performances of the banking system and the overall economy in the period preceding the signal horizon and the period after the signal horizon, where the period after the signal horizon is considered to be the crisis period. Division in these two periods remains due to the still existing crisis in Montenegro, therefore there is probability that results will be biased to some extent.

Babecký et al. (2012) emphasize that there are at least two problems with simple regression when there are many potential explanatory variables. First, putting all potential variables in one regression might significantly increase standard errors if irrelevant variables are included. Second, the use of sequential testing in order to exclude unimportant variables might lead to misleading results taking into consideration the fact that there is a probability that a relevant variable is excluded every time when the test is done. In order to solve these problems, the technique of model averaging is usually applied (Babecký et al., 2012, p.19). Bayesian model averaging considers model uncertainty by taking into account combinations of models and assigning them weights in accordance with their performance. There are only two papers related to the model uncertainty in the literature dealing with early warning systems, and one of them is related to systemic banking crises. Article by Crespo-Cuaresma & Slacik (2009) studied currency crises in 27 developing countries, and Babecký et al. (2012) studied banking, debt and currency crises in 40 developed countries.

The main limitation of early warning models for systemic banking crises in this paper is the fact that models are created on the basis of only one systemic banking crisis that happened in Montenegro. However, all requisite information cannot be provided by studying only one case. Taking into consideration the fact that not all banking crises happen according to the same pattern and when making conclusions just on the basis of a small number of events, there is a high probability that conclusions will be biased. Also, it is necessary to emphasize that in situations when an adequate database of historic data is available, general conclusions are often made on relative importance of individual indicators.

Selection of potential indicators is based mostly on the economic reasoning that takes into account theoretical assumptions and indicators already used in previous researches.

The choice of indicators also depends largely on the availability of data. Regarding the Montenegrin banking system, data on the monthly level are less available for the period until 2009, thus in terms of diversity it is more advisable to use quarterly data. However, concerning the data frequency, it is preferable to use monthly data because trends that indicate higher probability of crisis occurrence will be noticed earlier and necessary corrective measures will be undertaken in due time. Therefore, in this paper all indicators are used on the monthly basis starting from January 2005 to September 2012.

Variables in the paper which are not expressed as growth rates and interest rates, are expressed as natural logarithms. Applying the augmented Dickey-Fuller test for a unit root, it is determined that the most of time series are non-stationary. Therefore, non-stationary time series are differentiated, and by reapplying the ADF test after differencing time series it is determined that they are stationary. A few time series that are used in the paper have been differentiated two times in order to become stationary.

Although stationarizing implicitly brings the recent history of variables into the forecast, lagging of explanatory variables also allows varying amounts of recent history to be brought into the forecast. Therefore, lagging of explanatory variables enables predicting what will happen in the period t based on the knowledge of what happened up to the period t-1. A choice of the most adequate model is based upon the Information Criteria, what means that the model with the smallest value of the Schwarz Information Criterion (SIC) and the Akaike Information Criterion (AIC) is selected. Definitions of variables used in the paper are given in the following table.

Table 1: *Definitions of variables used in the paper*

Variable	Definition
ASSETS	Total assets at the aggregate level of the banking system
LOANS	Total gross loans at the aggregate level of the banking system
LLP	Total loan loss provisions at the aggregate level of the banking system
NET_LOANS	Total net loans at the aggregate level of the banking system, calculated as gross loans minus loan loss provisions
DEPOSITS	Total deposits at the aggregate level of the banking system
BORROWINGS	Borrowings from central banks, banks and other credit and financial institutions, and borrowings from the Government at the aggregate level of the banking system
CAPITAL	Total capital at the aggregate level of the banking system
LOANS_DEPOSITS	Loans-to-deposits coefficient at the aggregate level of the banking system
INT_INCOME	Total interest income at the aggregate level of the banking system
RESERVE_REQ	Total amount of reserve requirements at the level of the banking system
MONEX20	Index value that consists of twenty the most liquid companies on the Montenegrin stock exchange
PRICES	Annual growth rate of consumer prices in Montenegro
PRICES_M	Monthly growth rate of consumer prices in Montenegro
EURIBOR_1M	1-month EURIBOR
EURIBOR_3M	3-month EURIBOR
INDPR_SERBIA	Index of industrial production in Serbia
EUR_USD	Exchange rate EUR to USD

3 LOGIT APPROACH AND BAYES MODEL AVERAGING

As Wooldridge (2002; p. 530-533) suggests, considering models of binary response, their interest lies primarily in the response probability:

$$P(y = 1|x) = P(y = 1|x_1, x_2,..., x_k)$$

where *x* denotes a set of explanatory variables.

In order to avoid limitations of the linear probability model, it is necessary to consider the class of binary response models which have the following form:

$$P(y = 1|x) = G(\beta_0 + \beta_1 x_1 + ... + \beta_k x_k) = G(\beta_0 + x\beta)$$

where G is a function that takes values strictly between 0 and 1, for all real numbers *z*. This enables that estimated probabilities are strictly between 0 and 1. The expression xβ denotes $\beta_1 x_1 + ... + \beta_k x_k$.

The following expression should be considered[2]:

$$Pi = E(Y = 1|X_i) = \frac{1}{1 + e^{-(\beta_1 + \beta_2 X_i)}}$$

which may be denoted in a simpler way:

$$Pi = \frac{1}{1 + e^{-Z_i}} = \frac{e^Z}{1 + e^Z}$$

as a cumulative logistic distribution function, where $Z_i = \beta_1 + \beta_2 X_i$.

It is easy to verify that as Z_i ranges from -∞ do +∞, P_i ranges between 0 and 1, and that P_i is nonlinearly related to Z_i (i.e., X_i)[3]. P_i is nonlinear not only in X, but also in the parameters, what means that the standard OLS method can not be used for estimation of the logit model. However, this problem might be resolved in a relatively simple manner.

If P_i denotes the probability that crisis is going to occur, therefore $1-P_i$ denotes the probability that the crisis is not going to occur, presented like this:

$$1 - P_i = \frac{1}{1 + e^{Z_i}}.$$

[2] See: Gujarati, 2004; p. 595-597.

[3] *Gujarati* (2004, p. 595) notes that as $Z_i \to +\infty$, e^{-Z_i} tends to zero, and as $Z_i \to -\infty$, e^{-Z_i} increases indefinitely.

The previous expression may also be denoted as:

$$\frac{P_i}{1-P_i} = \frac{1+e^{Z_i}}{1+e^{-Z_i}} = e^{Z_i}.$$

The expression $P_i/1-P_i$ represents the odds ratio in favor of crisis occurrence – the ratio of the probability that crisis will occur to the probability that crisis will not occur. Therefore, if for example $P_i = 0{,}8$, it means that odds are 4 to 1 in favor of crisis occurrence.

If we take the natural log of the previous expression, we obtain:

$$L_i = \ln\left(\frac{P_i}{1-P_i}\right) = Zi = \beta_1 + \beta_2 X_i$$

where L, the log of the odds ratio, is not only linear in X, but also linear in the parameters[4]. L is called the logit model.

The criterion commonly used for determining the starting date of systemic banking crises is a share of nonperforming loans in total loans at the level of a banking system. Considering the threshold of a 10% share of nonperforming loans in total loans that is proposed by Demirgüç-Kunt & Detragiache (1998), the beginning of the systemic banking crisis in Montenegro should be June 2009 when this indicator reached 10.03%. However, few months earlier, deposits were withdrawn after a longer period of growth, and in the fourth quarter 2008 deposits decreased by -14.42% in comparison with the previous quarter. Furthermore, at the end of 2007, the Central Bank of Montenegro introduced a temporary measure of credit growth restriction since credit activity of banks has already become exaggerated. In accordance with the aforesaid, the author of this paper has determined October 2008 as the starting month of the crisis, when signs of crisis have already been shown in the form of deposit outflows. The signal horizon is defined 24 months prior to the crisis, what means that the dependent variable *y* takes the value 1 from November 2006 to October 2008. Estimation results of the dynamic logit model are presented in the following table.

Regarding nonlinear models, marginal effects give more information than coefficients. If only coefficients are taken into consideration, the size of change in probability of systemic banking crisis occurrence cannot be determined. Coefficients in the logit model show only the direction of change in probability, thus it shall be necessary to calculate marginal effects. Marginal effects of explanatory variables on dependent variable are presented in the following table.

[4] *Gujarati* (2004, p. 596) emphasizes that linearity assumption of OLS does not require that explanatory variables are linear, however linearity in the parameters is crucial.

Table 2: *Estimation results of the dynamic logit model*

Variable	Coefficient	Std. Error	z-Statistic	Prob.
C	-4.354124	1.189709	-3.659823	0.0003
LOANS	65.16109	20.44709	3.186815	0.0014
DEPOSITS	-45.13485	16.03267	-2.815181	0.0049
EURIBOR_1M	7.367738	2.893002	2.546745	0.0109
INDPR_SERBIA	-0.104783	0.050407	-2.078739	0.0376
LLP	31.47855	11.20501	2.809327	0.0050
EUR_USD	-23.04270	12.33094	-1.868689	0.0617
CAPITAL	26.51234	12.08045	2.194648	0.0282
LOANS_DEPOSITS_1	0.381331	0.167479	2.276891	0.0228
PRICES_3	1.180913	0.657362	1.796442	0.0724
McFadden R-squared	0.594652		Mean dependent var	0.269663
S.D. dependent var	0.446299		S.E. of regression	0.295038
Akaike info criterion	0.697293		Sum squared resid	6.876770
Schwarz criterion	0.976915		Log likelihood	-21.02953
Hannan-Quinn criter.	0.810000		Restr. log likelihood	-51.88017
LR statistic	61.70129		Avg. log likelihood	-0.236287
Prob(LR statistic)	0.000000			
Obs with Dep=0	65		Total obs	89
Obs with Dep=1	24			

Source: Author's calculations in EViews 6

Table 3: *Marginal effects*

Variable	Marginal effects
C	-0.206890
LOANS	3.096187
DEPOSITS	-2.144623
EURIBOR_1M	0.350085
INDPR_SERBIA	-0.004979
LLP	1.495731
EUR_USD	-1.094894
CAPITAL	1.259757
LOANS_DEPOSITS_1	0.018119
PRICES_3	0.056112

Source: Author's calculations in EViews 6

It is necessary to evaluate the predictive power of the estimated model. The cut-off value that separates the pre-crisis period from the normal period has been set at 0.5. The model has correctly predicted 88.76% observations. Furthermore, the model has precisely predicted the crisis in 79.17% cases (i.e. months), and the normal period in 92.31% cases. The model has proved to be unsuccessful in 11.24% cases. Prediction ability of the estimated logit model is presented in the following table.

Table 4: *Prediction ability of the estimated logit model*

	Estimated Equation			Constant Probability		
	Dep=0	Dep=1	Total	Dep=0	Dep=1	Total
P(Dep=1)<=C	60	5	65	65	24	89
P(Dep=1)>C	5	19	24	0	0	0
Total	65	24	**89**	65	24	89
Correct	**60**	**19**	**79**	65	0	65
% Correct	92.31	79.17	88.76	100.00	0.00	73.03
% Incorrect	7.69	20.83	11.24	0.00	100.00	26.97
Total Gain*	-7.69	79.17	15.73			
Percent Gain**	NA	79.17	58.33			
	Estimated Equation			**Constant Probability**		
	Dep=0	**Dep=1**	**Total**	**Dep=0**	**Dep=1**	**Total**
E(# of Dep=0)	58.26	6.74	65.00	47.47	17.53	65.00
E(# of Dep=1)	6.74	17.26	24.00	17.53	6.47	24.00
Total	65.00	24.00	89.00	65.00	24.00	89.00
Correct	58.26	17.26	75.52	47.47	6.47	53.94
% Correct	89.63	71.91	84.85	73.03	26.97	60.61
% Incorrect	10.37	28.09	15.15	26.97	73.03	39.39
Total Gain*	16.59	44.94	24.24			
Percent Gain**	61.53	61.53	61.53			

*Change in "% Correct" from default (constant probability) specification
**Percent of incorrect (default) prediction corrected by equation
Source: Author's calculations in EViews 6

Results of the Hosmer-Lemeshow test and the Andrews test are presented in the following table. A high value of the Andrews goodness-of-fit test and a low level of the Hosmer-Lemeshow test are desirable. Considering the Hosmer-Lemeshow test, if the associated p-value is significant ($p<0.05$), it might be an indication that the model doesn't fit the data. Since the H-L goodness-of-fit test statistic is much greater than 0.05, the null hypothesis that there is no difference between the observed and model-predicted values of the dependent variable is not rejected, implying that the model's estimates fit the data at an acceptable level.

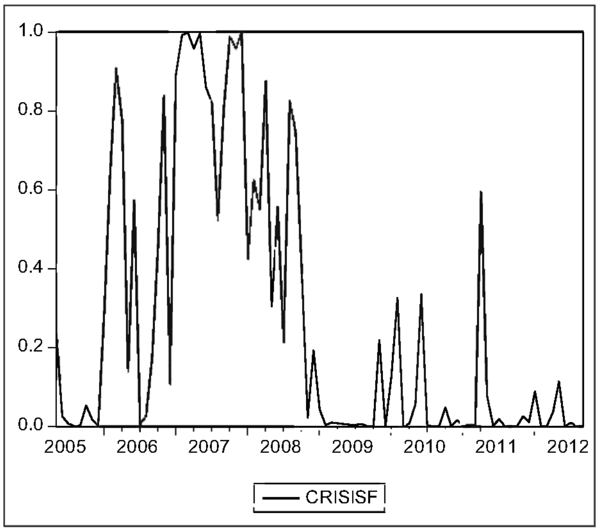
The next graph represents the forecasted probability of systemic banking crisis calculated from the dynamic logit model. The model sends signals within the signal horizon that is defined 24 months preceding the crisis – from November 2006 to October 2008. As it can be concluded from the graph, the highest probability of systemic banking crisis is during the first year of the signal horizon. This suggests that the model sends warning signals in the early stage, namely a year before the beginning of the crisis.

Table 5: *Results of the Hosmer-Lemeshow test and the Andrews test*

	Quantile of Risk		Dep=0		Dep=1		Total	H-L
	Low	High	Actual	Expect	Actual	Expect	Obs	Value
1	8.E-07	0.0002	8	7.99958	0	0.00042	8	0.00042
2	0.0004	0.0013	9	8.99276	0	0.00724	9	0.00725
3	0.0014	0.0043	9	8.97551	0	0.02449	9	0.02456
4	0.0049	0.0106	9	8.93227	0	0.06773	9	0.06824
5	0.0160	0.0423	9	8.77340	0	0.22660	9	0.23245
6	0.0479	0.1375	8	8.19409	1	0.80591	9	0.05134
7	0.1727	0.3344	7	6.71711	2	2.28289	9	0.04697
8	0.4049	0.6251	3	4.29662	6	4.70338	9	0.74874
9	0.6448	0.8755	2	1.80876	7	7.19124	9	0.02531
10	0.8926	0.9997	1	0.30990	8	8.69010	9	1.59156
		Total	65	65.0000	24	24.0000	89	2.79683
H-L Statistic			2.7968		Prob. Chi-Sq(8)		0.9465	
Andrews Statistic			42.1494		Prob. Chi-Sq(10)		0.0000	

Source: Author's calculations in EViews 6

Graph 1: *The forecasted probability of systemic banking crisis*



Source: Author's calculations in EViews 6

In order to check the robustness of obtained results, Bayesian model averaging is also apphed. As Babecký et al. (2012, p. 19-20) suggest, the following hnear regression model should be considered:

$$y = \alpha_y + X_y\beta_y + \varepsilon \qquad \varepsilon \sim (0, \delta^2 I)$$

where *y* is a dummy variable denoting crisis, α_y is a constant, β_y is a vector of coefficients, and ε is a white noise error term. X_y represents a subset of all available relevant explanatory variables, i.e. potential early warning indicators X. The number K of potential explanatory variables yields 2^K potential models. Mark γ is used to refer to one specific model from 2^K models. The information contained in models is then averaged using the *posterior* model probabilities that are considered under the Bayes' theorem:

$$p\ (M_{\gamma}, \mid y, X) \propto p\ (\ y \mid M_{\gamma}, X)\ p\ (M_{\gamma})$$

where $p(M_{\gamma}, \mid y, X)$ represents *posterior* model probability, which is proportional to the marginal likelihood of the model $p\ (y \mid M_{\gamma}, X)$ times the prior probability of the model $p\ (M_{\gamma})$.

The essence of Bayesian model averaging is assigning weights to estimated models in order to determine which models have the best performance. For this purpose it is necessary to calculate the Schwarz Information Criterion as one of the most commonly used information criteria in order to determine which specification is more appropriate for the data nature. This criterion is known as the Bayes Information Criterion which is actually approximation of the Bayes Factor. A higher value of weight is given to the model with a smaller value of SIC, thus the model that has a smaller value of SIC is considered to be a more favorable specification.

As already mentioned, using logit regression it is not possible to rank indicators according to their relative prognostic power when predicting systemic banking crises. This disadvantage can be partially overcome using Bayesian model averaging, because it is possible, by applying this technique, to assign adequate weights to simple logit models with at most two explanatory variables. Although individual variables do not have weights, their relative importance can be approximately determined on the basis of weights assigned to the model that contains these variables. Estimation results of implementation of the Bayesian model averaging technique are presented in the following table.

On the basis of weights assigned to individual models that are calculated using SIC, it may be concluded that estimated models have very similar performances. The best performance is that of the model with explanatory variables Monex20 which represents one of two indices on the Montenegrin stock exchange and net loans with weight 0.16408. The model with the lowest performances is one that contains variables - 3-month Euribor and monthly growth rate of consumer prices with weight 0.12907. Marginal effects of explanatory variables are presented in the following table.

Table 6: *Estimation results of implementation of the Bayesian model averaging technique*

Model	Variable	Coefficient	Statistic significance	Weight (0-1)
	ASSETS	106.23	0.0001	
Model 1	DEPOSITS	-69.62	0.0010	0.14370
	CAPITAL	13.42	0.0153	
Model 2	BORROWINGS	19.33	0.0003	0.13973
	LOANS	50.23	0.0000	
Model 3	RESERVE_REQ	-11.66	0.0205	0.15971
	EURIBOR_1M	5.35	0.0043	
Model 4	LLP	16.08	0.0024	0.13106
	LOANS_DEPOSITS	37.15	0.0010	
Model 5	INT_INCOME	7.60	0.0226	0.13266
	EURIBOR_3M	6.06	0.0138	
Model 6	PRICES_M	1.44	0.0113	0.12907
	MONEX20	-9.46	0.0011	
Model 7	NET_LOANS	47.32	0.0000	0.16408

Source: Author's calculations in EViews 6

Table 7: *Marginal effects*

Variable	Marginal effects
ASSETS	16.28
DEPOSITS	-10.67
CAPITAL	2.22
BORROWINGS	3.19
LOANS	7.46
RESERVE_REQ	-1.73
EURIBOR_1M	0.80
LLP	2.41
LOANS_DEPOSITS	5.87
INT_INCOME	1.20
EURIBOR_3M	0.98
PRICES_M	0.23
MONEX20	-1.25
NET_LOANS	6.24

Source: Author's calculations in EViews 6

Application of the Bayesian model averaging technique represents an important part of the analysis. Namely, this technique enables estimation of more variables that can be relevant indicators of systemic banking crises, than it would be possible by using only a regular logit model. Putting a higher number of variables in one single regression may cause problems, such as multicolinearity. As can be seen, the dynamic model has captured eight variables, while using the Bayesian model averaging technique 14 variables are included where six of them are the same as in the dynamic logit model. Instead of estimating only a set of simple logit regressions, Bayesian model averaging gives an insight into relative importance of some variables in comparison with other variables. Therefore, it is possible to determine which indicators are more reliable for prediction of systemic banking crises.

4 INTERPRETATION AND DISCUSSION

McFadden R^2 indicates a relatively good goodness-of-fit of the estimated model. Results of the estimated dynamic logit model suggest that loans have the highest marginal effect on the dependent variable. Therefore, if this indicator increases by 1%, the estimated probability of occurrence of the systemic banking crisis will increase by 3.10, holding constant the remaining variables. If the value of variable LLP that represents loan loss provisions increases by 1%, the probability of systemic banking crisis will increase by 1.50. Also, if the loans-to-deposits coefficient increases by 1%, the probability of systemic banking crisis will go up by 0.02. On the other hand, if deposits increase by 1%, the probability of systemic banking crisis will decrease by 2.14. If capital increases by 1%, the probability of systemic banking crisis will increase by 1.26.

Considering macroeconomic variables, it can be concluded that if 1-month Euribor increases by 1%, the probability of systemic banking crisis will go up by 0.35. Similarly, if EUR/USD exchange rate increases by 1%, the estimated probability of occurrence of systemic banking crisis will decrease by 1.09. Montenegro is a euroised economy, and one of the main advantages of fixed exchange rate regimes is that they enable achieving the macroeconomic stability thanks to a solid nominal anchor. However, it is necessary to emphasize that fixed exchange rates do not *a priori* provide macroeconomic stability. The main deficiency of fixed exchange rates is that they reduce flexibility of monetary policy. The reason for considering EUR/USD exchange rate as an early warning indicator is that Montenegro is a small and open euroised economy, so the trend of this variable might have a significant impact on the domestic economy. Concerning inflation, if the annual growth rate of consumer prices in Montenegro increases by 1%, the probability of systemic banking crisis will increase by 0.06.

One of the most important variables that are related to international indicators is economic growth of the country that represents the main trading partner of the domestic country. According to available data starting from 2005, the largest portion of Montenegro's trading exchange, taking into account both export and import, has been realized with Serbia, therefore the most significant trading partner of Montenegro is Serbia. If the index of industrial production in Serbia increases by 1%, the probability of systemic banking crisis occurrence will decrease by 0.005. It can be concluded that it is a variable with the lowest marginal effect in this model.

Seven simple logit regressions that individually have two explanatory variables are estimated, and thus there are 14 statistically significant indicators, while in the previous dynamic logit regression there are 9 indicators. Adequate weights have been assigned to all seven regressions using the technique of Bayesian model averaging. These results largely coincide with results of the previously estimated logit model.

If indicator that represents total assets in the banking system increases by 1%, the probability of systemic banking crisis occurrence will increase by 16.28, holding constant the remaining variables. Similarly, if loans increase by 1%, the probability of systemic banking crisis occurrence will go up by 7.46, and if net loans increase by 1%, the probability of systemic banking crisis occurrence will increase by 6.24. If loan loss provisions increase by 1%, the probability of systemic banking crisis occurrence will go up by 2.41. That can be explained by the fact that banks approved more risky loans during credit expansion, therefore, relatively shortly after that, they had to allocate a larger amount of loan loss provisions.

If deposits increase by 1%, the probability of systemic banking crisis occurrence will decrease by 10.67. Also, if the loans-to-deposits coefficient increases by 1%, the estimated probability that the systemic banking crisis will occur increases by 5.87. If capital increases by 1%, the probability of systemic banking crisis occurrence will go up by 2.22. Also, if borrowings which banks mostly take from their parent bank increase by 1%, the probability of systemic banking crisis occurrence will increase by 3.19.

Variable reserve requirements represent one of very few monetary instruments which the Central Bank of Montenegro has at its disposal, since Montenegro is a euroized economy. Actually, it is more appropriate to say that it is a liquidity instrument. If this variable increases by 1%, the probability of systemic banking crisis occurrence will decrease by 1.73. If 1-month Euribor increases by 1%, the estimated probability that systemic banking crisis will occur increases by 0.80, and with the increase of 3-month Euribor by 1%, the probability of systemic banking crisis occurrence will increase by 0.98.

If interest income increases by 1%, the probability of systemic banking crisis occurrence will go up by 1.20. Also, if the monthly growth rate of consumer prices in Montenegro increases by 1%, the probability of systemic banking crisis occurrence will increase by 0.23. Finally, if variable Monex20 increases by 1%, the probability of systemic banking crisis occurrence will decrease by 1.25.

Indicators relating to a credit boom thanks to very good performances, have a dominant role in early warning models for systemic banking crises. The accelerated economic growth influenced the banks to initiate the exaggerated lending activity that led to credit expansion with three-digit yearly credit growth rates, and that in turn even additionally encouraged overheating of the economy. Funds taken as borrowings from parent banks during the credit expansion were mostly used for the lending activity. It was just a question of time when it would come to the bursting of the bubble that reached enormous proportions especially on the housing market. Besides developments in the domestic banking sector and in the overall economy, the crisis occurrence is also accelerated by negative global trends influenced by the global economic crisis.

It is interesting that some indicators related to macroeconomic developments in the region and in the European Union, have also shown very good performances. These are 1-month Euribor and 3-month Euribor, EUR/USD exchange rate and the index of industrial production in Serbia. Therefore, it can be concluded that the Montenegrin economy and the banking system are exposed significantly to the trends on the global level. Developments on international markets have a significant impact on the domestic banking system and its stability, and therefore on the probability of systemic banking crisis occurrence.

5 CONCLUDING REMARKS

Although many economists, especially critics of economics as science, consider that these models have proved to be unsuccessful because they failed to predict occurrence of the present global crisis, the economic policy can not be conducted today in an appropriate and efficient manner without reliable quantitative information. However, it is necessary to take into account qualitative estimates made by economic experts. The use of early warning models for systemic banking crises have to be adequately integrated within broader analyses that take into consideration all important aspects, as it is inevitable that some of these aspects will be overlooked by one of these models. These models can have an important complementary role as an objective measure of the banking system vulnerability.

Regarding developing countries, it should be taken into account that they usually go through the catching-up phase in order to reach developed economies, and therefore they have higher economic growth rates. Economic growth during that phase is relying largely on the lending activity and it is sometimes difficult to differentiate between the credit expansion and the increased credit activity.

Results of the estimated models have shown that the systemic banking crisis in Montenegro has its roots in the domestic economy. Causes of crises originate from the period of unsustainable credit expansion. A very low level of credit activity during the period before the beginning of the credit expansion has encouraged banks to race for a market share. Also, results have shown that although roots of crisis are in the domestic economy, there is a significant impact of international trends on the Montenegrin banking system and overall economy.

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SHAREHOLDERS' PAY-OUT-RELATED THRESHOLDS AND EARNINGS MANAGEMENT

JERNEJ KOREN[1]
ALJOŠA VALENTINČIČ[2]

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ABSTRACT: *We investigate the thresholds in net shareholder pay-outs (dividends, share buy-backs and issuances) of a large sample of UK quoted firms. Discretionary accruals are analysed at these thresholds in relation to earnings management. By examining distributions and using a robust test for discontinuities, we show the existence of thresholds at zero bins of variables in question. Additionally, by looking at differences in means and medians of discretionary accruals in sorted distributions, we find that they are statistically different from bin to bin in vicinity of previously identified thresholds.*

1. INTRODUCTION AND PRIOR RESEARCH

Earnings, as the primary performance indicator of a firm, can be managed with the intent of companies reaching expectations-set performance thresholds (Burgstahler & Dichev, 1997) meeting analyst forecasts (Degeorge, Patel, & Zeckhauser, 1999), satisfying certain contractual obligations or fulfilling liabilities stemming from borrowing activities. Earnings management is also observed around certain corporate events, for example stock offerings or acquisitions (Erickson & Wang, 1999) or in connection with managers' compensations and bonus schemes (Bergstresser & Philippon, 2006). Still, earnings management cannot only be seen in a negative light. Under certain conditions it may also be beneficial for owners – through application of manager's acquired expertise in forecasting earnings or not dismissing a hired manager (who is good-working) too fast (Arya, Glover, & Sunder, 1998) – or at least neutral in a way that decisions taken with managed earnings in consideration are the same as they would be had earnings not been managed (Ronen & Yaari, 2010).

Among factors, assuming managers' threshold reasoning and, consequently, the possible appearance of earnings management, is also a company's dividend policy. Dividend policy is determined by the company's management and, as there is no unique rule about the dividend policy, similarly efficient and successful companies can – and do – have different dividend pay-outs (Brigham & Ehrhardt, 2005). Miller & Modigliani (1961) proposed

[1] University of Ljubljana, Faculty of Economics, Slovenia, Ljubljana, e-mail: jernej.koren@ef.uni-lj.si
[2] University of Ljubljana, Faculty of Economics, Slovenia, Ljubljana, e-mail: aljosa.valentincic@ef.uni-lj.si

a model of dividend irrelevance where corporate value should not be related to pay-out policy in a perfect and frictionless capital market.[3] Excluding taxes and transaction costs, investors should thus be indifferent between (cash) dividends and capital gains.

Historically, this has not been the case. Lintner's (1956) first study of dividend policy found that managers are reluctant to cut dividends and are willing to increase them only gradually after they are convinced of enough support of a higher level of dividends in the form of higher future earnings. Existing dividend levels thus act as a strong benchmark. In seeking to explain investor preferences for (cash) dividends, Shefrin & Statman (1984) put forward two explanations. Firstly, one of "self-control" where investors decide to consume only from dividends, not portfolio capital and are thus demanding dividends. Secondly, following Kahneman & Tversky's (1979) behaviour theory proposition that losses loom larger than gains, dividends are preferred by people who are averse to regret (a potential increase in share price had they sold their stock instead of receiving a dividend). The behaviourist view can also be a potential explanation for dividend decreases having a more negative market effect than dividend increases. If dividends and their levels present a benchmark for investors, market reactions to dividend changes, especially downward, are found to be substantial (e.g., Grullon, Michaely, & Swaminathan, 2002). Bhattacharyya (2007, pp. 9-10), for example, also provides a short overview of stylized facts on dividends.

A company's dividend policy can be affected by various factors such as market imperfections, behavioural considerations, firm characteristics or managerial preferences (Baker, Powell, & Veit, 2002). They differ in importance to individual firms, but they form the basis for possible earnings management. While the latter two factors include firm- and management-specific factors, the former two factors comprise broader aspects such as different tax treatment of dividends and capital gains, overcoming information asymmetries with signalling new or additional information and shareholder and investor clienteles that favour dividends in various degrees at various times (see Baker & Wurgler (2004) for a catering theory of dividends).

The distributional analysis and existence of thresholds was first suggested by Hayn (1995) who points out the discontinuity of earnings around zero in her study of the information content of loses.[4] Building on this empirical irregularity, Burgstahler & Dichev (1997) show that firms manage earnings to avoid reporting loses or earnings decreases. They interpret low frequencies of small loss (earnings decline) observations and high frequencies of small profit (earnings increase) observations as a consequence of firms' active efforts to cross the loss (earnings decline) threshold what results in a migration of observations to the right of such divide as seen if a distribution is plotted. Assuming that without

[3] DeAngelo & DeAngelo (2006) contested that pay-out policy is not irrelevant as put forward by Miller & Modigliani (1961) but their proposition was reconciled as having assumed different agency costs (Handley, 2008).

[4] An interesting case of goal reaching behaviour research is also the analysis of Carslaw (1988) who finds abnormal distribution of income numbers in financial statements with the bias tilting towards numbers just above multiples of powers of ten (i.e., $N \times 10^k$) as cognitive reference points.

earnings management the distribution of earnings would be fairly smooth, they test the documented asymmetry around zero (earnings or changes in earnings) thresholds.

Their findings are confirmed by Degeorge, Patel, & Zeckhauser (1999) who add another threshold of meeting analyst forecasts (i.e., avoid earnings surprises). Additionally, they establish a hierarchical order of the three with positive earnings threshold being predominant, followed by not falling short of previous earnings and lastly meeting analyst expectations. Critique of distribution analysis is based mainly on the effect of deflator and the sample selection procedure, both of which can have an effect on the resulting distribution (Durtschi & Easton, 2005). If the deflator differs systematically between profit and loss firms it can move the scaled observations towards or away from zero, what is most commonly the case when scaling by market price, but also found for other deflators (Durtschi & Easton, 2009). Alternative explanations of the discontinuity include asymmetric effects of taxes and special items that also contribute to observed shapes of distributions (Beaver, McNichols, & Nelson, 2007).

We therefore study thresholds and earnings management from the standpoint of attaining (expected) pay-outs to investors as earnings levels are often directly or at least indirectly connected to the pay-outs, e.g., in companies with fixed pay-out ratio policy or linked to various contractual obligations that set limits on pay-out possibilities. The first study in this area is the analysis of Finnish companies that managed earnings to ensure constant dividend pay-out to large institutional investors who prefer stable dividends (Kasanen, Kinnunen, & Niksanen, 1996), whereas, in the US, Daniel, Denis & Naveen (2008) have shown that firms manage earnings upward to reach expected levels of dividends (defined as last year's dividend) when they expect they would otherwise fall short of it, proving they are important thresholds for managers. Similar findings are reported by Atieh & Hussain (2012) for UK. They show that earnings may be managed by firms which also try to avoid a decrease or even elimination of dividends and show a concern for coverage ratios, but the pressure is lower for larger firms which face less restrictive debt covenants. Debt covenants can impose restrictions on dividend payments if the financial position of the firm does not appear adequate. Moir & Sudarsanam (2007) report three quarters of firms in their study to have covenants attached to debt contracts. Another recent study by Bennet & Bradbury (2007) proposes dividend cover to be considered as a threshold as firms are likely to manage earnings to avoid cutting dividends, i.e., keeping them at least at their prior year's values.

A comprehensive survey of CFOs by Brav et al. (2005) shows that managers are willing to go great lengths to avoid a dividend cut but increases in dividends are a second-order concern. The authors also observe that share repurchases have become an established alternative pay-out instrument to dividends. However, they do not convey the same signals about companies' future behaviour or performance. Dividends are seen as a more permanent commitment to provide shareholders with a reasonably stable cash flow, whereas repurchases (particularly the ones on a discretionary and non-constant basis) are viewed as more flexible. Repurchases would now be the primary choice of many firms had their dividend history not existed. Interestingly, little support is found for the signalling hypotheses, that is, not many managers state they are paying dividends to convey a company's true state (future prospects) or to intentionally separate them from competitors. Taxes are also not a primary concern in deciding about the payment/increase of dividend or in choosing between them and repurchases.

Repurchases are gradually replacing dividends as the primary pay-out method with higher correlation to possible swings in earnings levels (with a shorter lag than for dividends). Skinner (2008) reports that firms which pay dividends only practically do not exist anymore. Other research has also found a decline in dividends paid by US listed firms, attributing it to both different firm characteristics and lower propensity to pay in general (Fama & French, 2001). Contrary to the latter, Grullon & Michaely (2002) find repurchases to be important in substituting dividends and US corporations financing them with funds that would have been otherwise used for dividend increases. What further motivates our research is a finding of Hribar, Jenkins & Johnson (2006) who assert that share repurchases are used by some companies to reach analysts' earnings per share forecasts. This implies that repurchases might be viewed as a possible earnings management tool.

In this paper we analyse a UK sample with focus on three theoretically possible shareholder-related cash flows.[5] Next to dividend pay-outs we also consider share repurchases and issuances of new shares, where the company is receiving funds from investors, resulting in a "negative" pay-out to shareholders. As these three shareholder-related cash flows might all be broadly regarded as dividend (pay-out) related decisions, we investigate the existence of thresholds in all three cases. This view is in line with Ohlson's (1995) valuation model that confirms Miller & Modigliani (1961) value displacement property as dividends are paid out of book value and consequently reduce market value on an one-for-one basis rendering dividend policy irrelevant. Ohlson's model allows (requires) negative net dividends, i.e., capital contributions (share issuances) exceeding pay-outs.

As accruals, and more precisely their discretionary component, are often associated with lower earnings quality and possible earnings management, (e.g., see Dechow, Ge & Schrand (2010) for an overview) we are also interested to what extent discretionary accruals are present at the hypothesized pay-out thresholds. Although Yong & Miao (2011) find that dividend paying status is associated with the quality of earnings in general, they also find that the association is stronger when dividends increase in size. Therefore, inspecting the margin of dividend payment or dividend increase would be informative since firms potentially having difficulties in reaching these thresholds could still make use of discretionary accruals to arrive to them.

H1: Companies attempt to reach thresholds of net shareholders pay-outs, which results in breaks in distributions of net shareholder pay-outs.

H2: Thresholds are associated with significant discretionary accruals levels.

[5] Beginning of section 3 (*Sample selection and description*) explains our choice of the UK market.

This study helps to determine if repurchases and new share issuances, although not typically regular events, affect the pay-out level targeted by the management. This would broaden the perception of flows that are viewed as important in setting companies' dividend policy. In the process, a robust test of discontinuity of distribution is used (Garrod, Ratej Pirkovic, & Valentincic, 2006). Moreover, discretionary accruals as a proxy for earnings management are analysed in relation to the pay-out levels.

The remainder of the paper is structured as follows. Section 2 presents the research design employed in our analysis, followed by sample selection and data description in the next section. Section 4 presents main empirical results and section 5 reports additional tests. Section 6 concludes.

2. RESEARCH DESIGN

We begin our investigation by constructing the variables representative of pay-out-related thresholds. Typically, dividend pay-outs are investigated, either in their total amount or as change from year to year, both relative to opening total assets to account for size differences among firms. We denote DIV as the ratio of dividends to lagged total assets and D_DIV as the ratio of change in dividends from the previous year to the current year, scaled by lagged total assets. The variable D_DIV_DIV scales the dividend change from the previous year to the current year by previous year's dividends level to get a variable representing relative yearly pay-out changes.

We calculate net shareholder cash flows as the sum of all cash flows investors might be dealing with, i.e., dividends received plus stock repurchases (as positive cash flows from the company to shareholders) less any share issuances in a given year (negative cash flows from shareholders to the company):

net shareholder cash flows = dividends + share repurchases – share isuances

Analogous to the dividend variables above, NSCF denotes the ratio of net shareholder cash flows to lagged total assets, D_NSCF scales yearly changes in net shareholder cash flows by lagged total assets and D_NSCF_NSCF is the change in net shareholder cash flows scaled by its lagged value. We also calculate and perform initial analyses on the scaled sums of dividends and stock repurchases only but, as dividends are highly dominating this sum, the results do not differ in any important way from dividend-only findings and offer no incremental insights. This part is therefore not investigated further in this paper.

Variables as defined above are then distributed into bins of widths 0.005 for total assets scaling and 0.01 for pay-out scaling.[6] That corresponds to forming groups that contain

[6] These bin widths were selected for both, comparability with prior research investigating distributions, although of different analysed and scaling items (Burgstahler & Dichev, 1997; Durtschi & Easton, 2005; Bennet & Bradbury, 2007) and ease of interpretation. As setting the bin width can have a huge effect on the histogram

observations with values within 0.5% of lagged total assets or 1% of lagged pay-out. These increments are also used for all subsequent bins. Bin widths for pay-out scaling are twice as big as for total assets scaling because the latter are much larger in absolute value and their use as a denominator results in much smaller ratios that have to be presented with higher accuracy to prevent artificial "histogram over-smoothing" (Scott, 1979). All bins are defined to include lower bound as we want the central bin to include observations with zero and small positive values and exclude the upper bound. Although debatable, we consider zero (as the scaled amount, or change, where applicable) as the first non-negative signalling value and thus the threshold to investigate. The so-called "zero bin" is therefore defined as including x if $0 < x < 0.005$ or $0 < x < 0.01$ in the case of scaling the variables by lagged pay-out. We also draw attention to the distinction between the terms used in subsequent analysis. The terms central bin, zero bin and bin(0) all denote the first bin of the distributions immediately to the right of zero, containing smallest positive scaled observations and the exactly-zero observations at its lower bound. The bins further to the right are denoted as bin(1), bin(2), etc., and bins further to the left are denoted as bin(-1), bin(-2), etc. We use the expressions zero observations, zero values or zeros to denote observations in the central bin for which the values of the variables analysed exactly equal to zero.

We plot histograms for all variables in this paper, with and without zero values. We do so because we expect a large number of observations to have the value of zero (either no dividends are paid out or the dividend level is the same as in the previous year, resulting in zero change) and we investigate whether zero values are predominant in threshold-attaining behaviour or does this behaviour exist without zeroes as well. Furthermore, a visual inspection of the distributions might reveal other potential points (bins) of interest that would be investigated further.

To formally test our assumption that zero bins in dividends and net shareholder cash flows are a valid and valuable threshold for companies, we employ a robust test for discontinuity of a distribution proposed by Garrod, Ratej Pirkovic & Valentincic (2006) – henceforth GRPV test – which requires no strict assumptions about the underlying distribution that one is testing. Requiring an assumption of normality in the test statistic was a shortcoming that accompanied (dis)continuity of distribution tests applied thus far in the literature, e.g., in Burgstahler & Dichev (1997). Developed with earnings-management applications in mind, the GRPV test is especially reliable in samples with more than 5,000 observations, a number that we easily exceed in our analysis.

being constructed (Wand, 1997), we considered various alternative widths in the process. While the software suggested widths for histograms ofmtervals in the following section were between 0.0045 and 0.0048 for total assets scaling and between 0.010 and 0.011 for pay-out scalmg, various optimal bin width formulas (Scott, 1979; Wand, 1997; Garrod, Ratej Pirkovic & Valentincic, 2006) produced a span ofresults. These ranged from widths of below 0.001 using Freedman-Diaconis' formula ($h = \frac{2 \times IQR}{\sqrt[3]{n}}$) to over 100 using the Sturgers' rule ($h = \frac{x_{max} - x_{min}}{1+\log_2 n}$), also dependent on the variable. The latter widths were drastically reduced to 0.400 or less if outliers at 1% were removed before the calculation. Suggested bin widths obtained using the Scott's formula ($h = 3.49 \times \sigma n^{-1/1}$) were between the two extremes.

The GRPV test statistic τ, derived from Chebyshev inequality, is computed as follows in equation (1) below, while assuming independent events gives us the inputs $E(X_i) = N \times p_i$ and $(X_i) = N \times p_i \times (1 - p_i)$, where N is the total number of observations in the sample and p_i is the probability that an observation will fall into interval (i), primarily computed as an average of two adjacent intervals: $p_i = \frac{X_{i-1} - \tilde{X}_{i+1}}{2N}$.

$$\tau_i = \frac{\tilde{X}_i - E(X_i)}{\sqrt{var(X_i)}} \tag{1}$$

where $\tilde{X}_i$ is the actual number of observations in interval (i). Values of the test statistic are tabulated in Garrod, Ratej Pirkovic & Valentincic (2006) and a break in the distribution at interval i is identified at standard significance levels of 10%, 5% and 1% corresponding to absolute values of the τ statistics of 3.16, 4.47 and 10.00 respectively.

We are also interested in the role accruals have at the hypothesized thresholds and in particular the discretionary component of total accruals. We use the modified Jones model (Dechow, Sloan, & Sweeney, 1995) to estimate non-discretionary accruals which we then use to determine the discretionary component of accruals as the difference between estimated values and total accruals. Firstly, total accruals are computed as:

$$TACC = (\Delta CA_t - \Delta CL_t - \Delta Cash_t + \Delta STD_t - Dep_t)/TA_{t-1} \tag{2}$$

where ΔCA_t is the change in current assets, ΔCL_t is the change in current liabilities, $\Delta Cash_t$ is the change in cash and cash equivalents, ΔSTD_t is the change in short term debt, Dep_t are depreciation and amortization charges and TA_{t-1} are lagged total assets. The modified Jones model is of the following form:

$$NDACC = \alpha_1(1/TA_{t-1}) + \alpha_2(\Delta REV_t - \Delta REC_t) + \alpha_3(PPE_t) \tag{3}$$

where TA_{t-1} are lagged total assets, ΔREV_t is the change in revenues, scaled by lagged total assets, ΔREC_t is the change in receivables, scaled by lagged total assets and PPE_t is gross property plant and equipment, scaled by lagged total assets. Estimates of α_1, α_2 and α_3 are obtained by estimating the model in equation (3) by industries with total accruals on the left-hand side. The estimated coefficients are then used to generate non-discretionary accruals.

The residuals of this model are discretionary accruals. Discretionary accruals are then analysed bin-wise for possible differences in their mean or median values. For this purpose, the t-test for means and the Wileoxson rank-sum for medians are used. We expect statistically significant differences of discretionary accruals in bins around zero thresholds that would link the two potential indicators of earnings management and suggest discretionary accruals' use in connection with these thresholds.

3. SAMPLE SELECTION AND DESCRIPTION

We acquire data of publicly listed UK companies from *Datastream*. This market is selected because companies in the UK have historically paid considerable dividends that still persist. A large majority (almost 85%) of UK firms paid dividends in the 1990s and dividend pay-outs dominated proportion-wise, although repurchases have been on the rise (Renneboog & Trojanowski, 2005).[7] Even recently, despite the trend of declining pure dividend pay-outs (Skinner, 2008), UK firms still tend to pay out dividends relatively more often than elsewhere (Denis & Osobov, 2008). As we want to have the initial sample as broad as possible, companies in the period from 1990 to 2012 are considered. Prior to 1990, the lack of data availability hinders a more detailed analysis and an incomplete set of companies' financial information was provided for 2012 at the time of data collection.

A note is necessary about dividend inputs from the database. Since IFRS-compliant reporting became mandatory for all listed companies in the EU for annual periods beginning on or after 1st January 2005, a provision in the standards requires companies to account differently for dividends paid. Before 2005, under the Statement of standard accounting practice (SSAP 17 - Accounting for post balance sheet events, 1980), dividends were accounted for as an adjusting post balance sheet event in the period to which they related. After 2005, it is prohibited to recognise dividends declared after the end of reporting period as a liability to that same reporting period (IAS 10 - Events after the reporting period). Instead, such dividends are disclosed in the notes but accounted for in the period in which they are paid. Thus, actual pay-out liability has priority over its source (earnings). This results in reported dividends in period (t) consisting of final dividend for period (t-1) and interim dividend(s) for period (t). Final dividend for period (t) is then recognised in period (t+1) financial statements etc.[8]

Table 1: *Sample construction procedure*

firm-year observations of listed UK companies in the period 1990 – 2012	38,429
observations with missing essential data	742
observations with zero total assets or sales	2,065
observations of financial and utility firms	5,380
final sample firm-year observations (3,177 distinct firms)	30,242

Notes: This table presents sample selection process. Starting sample of listed UK companies is obtained from Datastream and identified using nation code (WC06027). All financial industry related firms and utilities are excluded due to their specific operating properties.

[7] Dividend payments have been more frequent in the UK also due to the more favourable tax treatment of dividends in the past (prior to the Finance Act 1997, see section *Additional tests* for more information) but remained high after the change as well.

[8] For example, GlaxoSmithKline (GSK) in its 2005 annual report shows a breakup of dividends into four interims of (all in £m) 568, 567, 568 and 792 for 2005 respectively and 575, 573, 571 and 684 for 2004 respectively. But, since GSK normally pays a dividend two quarters after the quarter it is relating to, dividends actually paid in 2005 were the last two interims for 2004 and the first two interims for 2005. The sum of those, £m 2390, is then reported as dividends for 2005 and also found as the database entry.

We first eliminate entries with missing data that are essential for the analysis, e.g., missing total assets or industry codes. We then remove observations with zero total assets and/or zero sales as these are not believed to be truly operational and the former would imply division by zero in the construction of our variables of interest. Lastly, as a common step, we remove firms from financial and utility sectors because of their operating specifics. We end up with 3,177 distinct companies and 30,242 firm-year observations as presented in Table 1. Out of these, 62% include dividend payments, 60% report proceeds from sale or issuance of stock and 11% show a change in redeemed, retired or converted stock. A substantial share of issuances indicates a possible large effect on NSCF, whereas the extent of repurchases is somewhat smaller than expected. Examination of the data also revealed some confounding entries in form of negative values of repurchases (14 identified) and negative values of issuances (134 such cases), both of them are not supposed to be negative following the definition of *Datastream* datatypes. A subset of each was, where possible, manually checked back against firms' annual reports and entries were corrected accordingly, e.g., into positive values. Lastly, otherwise sound observations with missing dividends, repurchase or issuance data had those set to zero.[9]

Table 2 presents sample structure by years. As there are no big deviations in any specific year, we can observe a first peak in the number of listed UK companies in 1997, followed by a slight decrease and another gradual but steady increase in the years following up to 2006. However, in the last years there is quite a strong decline coinciding with the development and deepening of the financial crisis. Data for 2012 were not fully populated at the time they were collected.

Table 2: *Year composition*

Year	N	in %	Year	N	in %	Year	N	in %
1990	1,154	3.82	1998	1,462	4.83	2006	1,551	5.13
1991	1,166	3.86	1999	1,389	4.59	2007	1,491	4.93
1992	1,147	3.79	2000	1,398	4.62	2008	1,352	4.47
1993	1,152	3.81	2001	1,443	4.77	2009	1,236	4.09
1994	1,184	3.92	2002	1,474	4.87	2010	1,124	3.72
1995	1,183	3.91	2003	1,502	4.97	2011	1,063	3.51
1996	1,467	4.85	2004	1,553	5.14	2012	658	2.18
1997	1,542	5.10	2005	1,551	5.13			

Notes: Year distribution of the sample in presented in this table. Total number of observations is 30,242. At the time of data collection year 2012 was not fully populated, therefore the number of observations is accordingly smaller.

Descriptive statistics in Table 3 suggests skewed distributions in almost all variables. As we are interested in the centre of distributions and especially in specific breaks, quartiles are reported along with the average, but standard deviations indicate that there are

[9] There were 1,521 such cases, of those only 390 with missing dividends. Remaining missing repurchases and/or issuances would prevent the construction of NSCF with dividends mostly available. Omission of these cases does not change the results.

substantial extreme observations.[10] The number of observations is mostly affected by the denominator, particularly when scaling by past dividends and less so when scaling by past NSCF. The first four variables use lagged total assets for scaling and are limited by that. Only DIV and D_DIV have comparable means and medians, dividends amounting on average to around 2% of previous year's total assets and dividend change being positive but of minor amount compared to total assets. The remaining four variables have means and medians differing in both sign and magnitude, once more implying skewed distributions.

Table 3: *Descriptive statistics*

Variable	Mean	25%	Median	75%	SD	N
DIV	0.024	0	0.015	0.033	0.086	26813
NSCF	-0.256	-0.001	0.009	0.030	6.274	26813
D_DIV	0.002	0	0	0.004	0.097	26813
D_NSCF	-0.156	-0.012	0.001	0.016	9.047	26813
D_DIV_DIV	0.425	-0.004	0.084	0.241	11.836	17201
D_NSCF_NSCF	39.208	-1	-0.039	0.229	2,175	22829

Notes: This table presents descriptive statistics for analysed variables. DIV = dividends (WC04551) scaled by lagged total assets (WC02999); NSCF = (dividends + repurchases (WC04751) – issuances (WC04251)) = net shareholder cash flows scaled by lagged total assets; D_DIV = change in dividends from year (t-1) to (t) scaled by lagged total assets; D_NSCF = change in net shareholder cash flows from year (t-1) to (t) scaled by lagged total assets; D_DIV_DIV = change in dividends from year (t-1) to (t) scaled by lagged dividends; D_NSCF_NSCF = change in net shareholder cash flows from year (t-1) to (t) scaled by lagged net shareholder cash flows. Number of observations varies due to availability of respective denominators used in variables' construction.

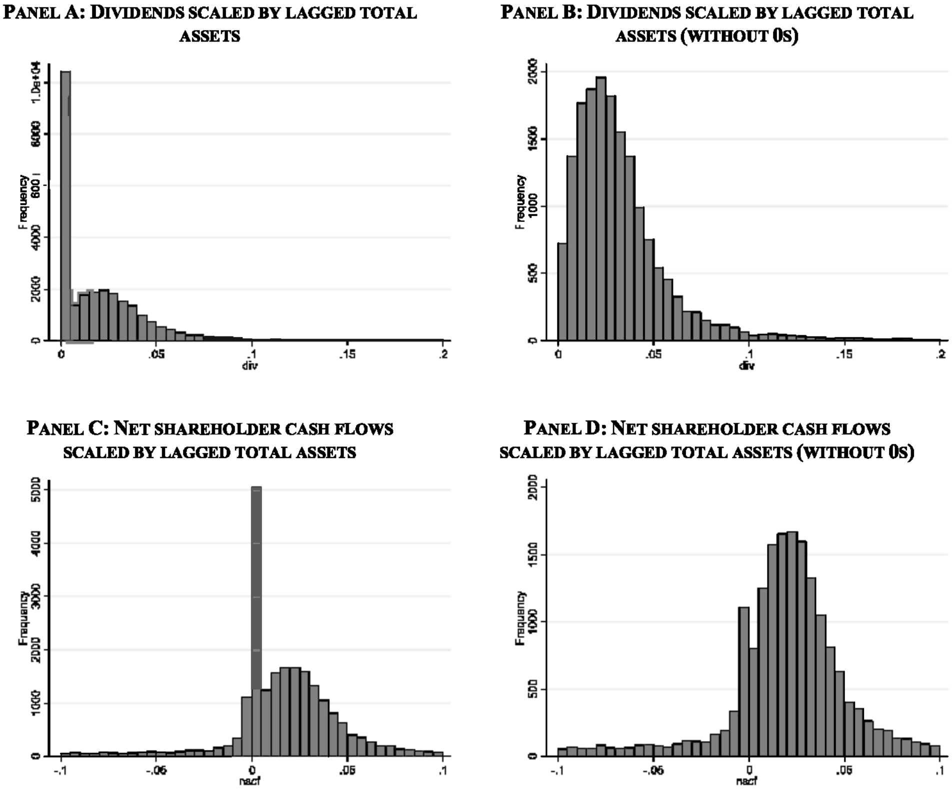
For visual representation, we plot histograms of respective variables sorted into bins 0.005 or 0.01 wide as described in the previous section to arrive at distributions of interest. Almost all distributions imply a threshold at the zero bin, firstly in amounts relative to total assets (attaining dividends or non-negative net shareholder cash flows). Panels A and C in Figure 1 show striking mode bins of small non-negative pay-outs and a comparison of the two panels suggests that dividends clearly dominate also in NSCF calculation. Although halved in size (10,419 observations in bin(0) for DIV and 5,047 observations in bin(0) for NSCF), the zero bin of the latter is still clearly outstanding from the remaining distribution. There are also changes, with observations shifted to bins left of zero due to effect of issuances, but the distribution to the right of zero is not much different compared to DIV.

Bin(0) modes disappear when observations equalling exactly zero are excluded in panels B and D. What remains is a mode in some of the subsequent positive bins (around 2-3% of lagged total assets) for both DIV and NSCF. While the zero bin in DIV is not standing

[10] We did not exclude any outliers since our central analysis is concerned with specific observations at the centre of respective distributions. As all our variables are ratios, outliers can arise due to disproportionate numerators and denominators in the span of one year. This may be related to one variable only. Therefore, by excluding outliers relating to one variable we could lose economically-sound observations in other variables.

out in any way, the one in NSCF is missing almost 400 observations (estimated as the difference to the average of adjacent bins) for a smooth, normal-like distribution. This case could indicate that NSCF are not a threshold of their own, in a way that firms would target its combined value as a reference point for investors.

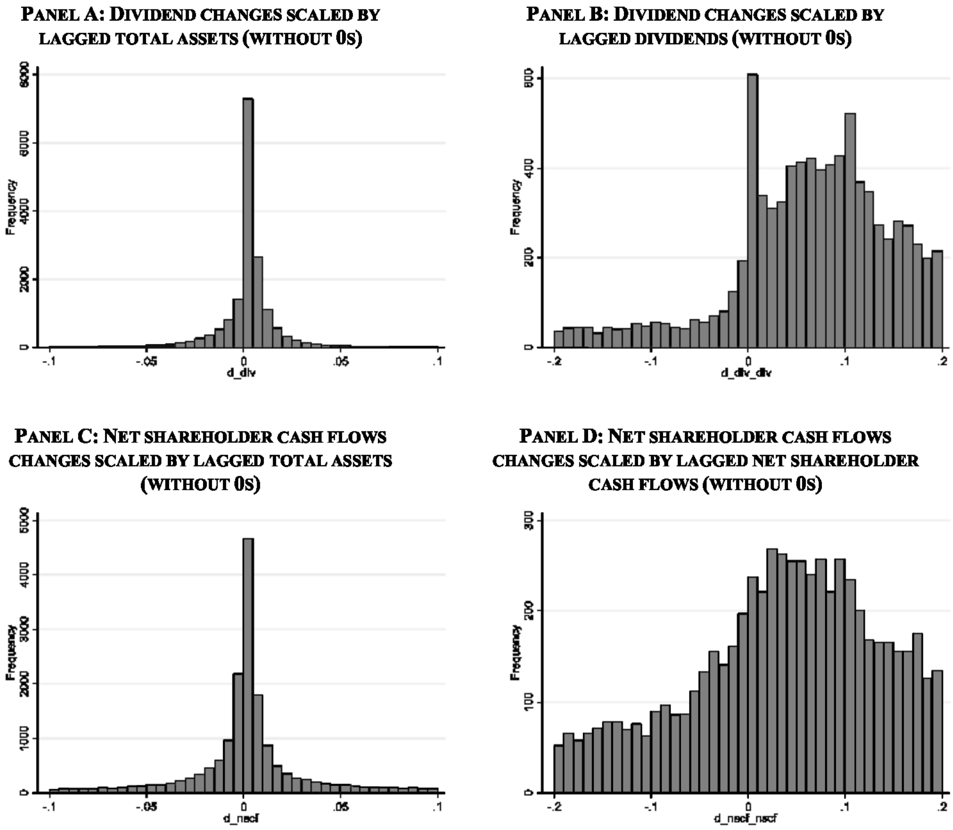
Figure 1: *Histograms of selected distributions*



Notes: This figure presents distributions of variables of interest. Panels A and B graph DIV, with and without zero observations, and panels C and D graph NSCF, with and without zero observations. DIV = dividends (WC04551) scaled by lagged total assets (WC02999) and NSCF = (dividends + repurchases (WC04751) – issuances (WC04251)) = net shareholder cash flows scaled by lagged total assets.
Bin width is 0.005 with lower bound inclusion, i.e., "zero bin" includes x if 0 < x < 0.005, "bin one" includes x if 0.005 < x < 0.01 etc.

As observations of zero in given variables have such an overwhelming effect on distributions, they are not reported in FIGURE 2 but they are still included in the analysis that follows. Findings of clearly modular bin(0) are confirmed for scaled changes in dividends (D_DIV, Panel A) and scaled changes in net shareholder cash flows (D_NSCF, Panel C) – even without observations equalling exactly zero. What is of interest is that, in case of D_NSCF, the bin with the second highest frequency is actually the first negative (and not positive, as more commonly expected) bin and this pattern is even repeated bin-wise as we move away from zero bin. The negative effect issuances have on D_NSCF outweighs the combined positive effect of dividends and repurchases in these cases.

Figure 2: *Histograms of selected distributions*



Notes: This figure presents distributions of variables of interest. Panel A graphs D_DIV, panel B graphs D_DIV_DIV, panel C graphs D_NSCF and panel D graphs D_NSCF_NSCF, all without zero observations. D_DIV = change in dividends (WC04551) from year (t-1) to (t) scaled by lagged total assets (WC02999); D_DIV_DIV = change in dividends from year (t-1) to (t) scaled by lagged dividends; D_NSCF = change in net shareholder cash flows (= dividends + repurchases (WC04751) – issuances (WC04251)) from year (t-1) to (t) scaled by lagged total assets; D_NSCF_NSCF = change in net shareholder cash flows from year (t-1) to (t) scaled by lagged net shareholder cash flows.
For panels A and C bin width is 0.005 with lower bound inclusion, i.e., "zero bin" includes x if $0 < x < 0.005$, "bin one" includes x if $0.005 < x < 0.01$ etc., and for panels B and D bin width is 0.01 with lower bound inclusion, i.e., "zero bin" includes x if $0 < x < 0.01$, "bin one" includes x if $0.01 < x < 0.02$ etc.

Lastly, looking at pay-out changes relative to their lagged values (D_DIV_DIV and D_NSCF_NSCF, Panels B and D, respectfully), zero bin threshold mode remains obvious in dividend changes scaled by lagged dividends, but with a lot lesser difference to surrounding bins. In the case of D_NSCF_NSCF zero bin practically blends in the distribution and does not even seem to represent a threshold on the left (negative) side, the distribution itself not displaying any noticeable breaks whatsoever. This is once more suggestive that no systematic threshold attaining behaviour can be observed with regard to net shareholder cash flows.

Frequencies of dividend increases and net shareholder cash flows increases relative to their lagged values rise and/or remain high up to bins denoting growth in the order of 10% (note that bin width is 0.01 in these two cases as the denominators are considerably smaller than total assets used beforehand). Another interesting observation is bin(10) of D_DIV_DIV, denoting cases of dividend increase between 10% and 11% compared to previous year's dividends. The bin in question appears to jut out of the distribution and is also statistically evaluated in the next section.

4. RESULTS

We attempt to formally confirm observations derived from histograms in the previous section with the use of GRPV discontinuity of distribution test. Table 4 reports values of the GRPV test applied for all cases inspected earlier (with and without zero observations) and fully confirms our assumptions. In all six cases of zero values of variables included, zero bin represents a discontinuity from the remaining distribution, inferences being done at P-values far below 1% (critical values of the test in absolute terms for significance levels of 10%, 5% and 1% are 3.16, 4.47 and 10, respectively). The discontinuity is stronger in dividend-related variables compared to NSCF-related ones, implying that repurchases and issuances lessen the break to some extent by moving some observations away from zero bin. Scaling by total assets results in stronger breaks than scaling by lagged values of pay-out, suggesting that the choice of scaling variable also plays an important role in distribution analysis as also suggested by previous research (Dechow, Richardson, & Tuna, 2003; Durtschi & Easton, 2005).

On the other hand, in cases where zero values of variables are excluded from distributions, discontinuity is still statistically confirmed in four out of six cases. The H_0 of continuity of distribution cannot be rejected in the first (DIV) and last case (D_NSCF_NSCF) as suggested and anticipated by the histograms in the preceding section, whereas other variables have results significant at the 1% level although test values are considerably lower than before the exclusion of zeros. A comparison of the four variables representing scaled changes in either dividends or net shareholder cash flows shows consistently larger breaks in dividends. We thus regard them as the driving factor for threshold existence. The fact that breaks are lessened with the inclusion of repurchases and new share issuances implies that these are not used with the intent of reaching a NSCF-related threshold, but rather for other purposes.

Table 4: *GRPV discontinuity of distribution test*

	(1) zero observations included	(2) without zero observations
DIV	376.61	1.42
NSCF	115.39	-11.36
D_DIV	336.51	123.70
D_NSCF	124.24	62.61
D_DIV_DIV	73.02	21.16
D_NSCF_NSCF	39.97	1.95

Notes: Reported are τ values of GRPV discontinuity of distribution test for zero bins of variables analysed. First column reports test statistics computed including observations of zero in selected variables, and second column reports test statistics computed without these observations.
With H_0: *the distribution function is continuous*, the values of τ at standard levels of significance are: at 10% $|\tau| = 3.16$, at 5% $|\tau| = 4.47$, and at 1% $|\tau| = 10$. As the number of observations in adjacent bins is required by the test, in the first row (case of DIV) bins on the left of zero (negative bins) are empty (there are no negative dividends) and are as such affecting test statistic computation.
DIV = dividends (WC04551) scaled by lagged total assets (WC02999); NSCF = (dividends + repurchases (WC04751) – issuances (WC04251)) = net shareholder cash flows scaled by lagged total assets; D_DIV = change in dividends from year (t-1) to (t) scaled by lagged total assets; D_NSCF = change in net shareholder cash flows from year (t-1) to (t) scaled by lagged total assets; D_DIV_DIV = change in dividends from year (t-1) to (t) scaled by lagged dividends; D_NSCF_NSCF = change in net shareholder cash flows from year (t-1) to (t) scaled by lagged net shareholder cash flows.

We therefore confirm breaks at zero bins in the distributions of scaled pay-outs, which is indicative of existence of thresholds. The exclusion of zero observations has different meanings, depending on the variable in question. The DIV variable is specific, as it is bounded to the left of zero, i.e., there are no negative cash dividends. Zero observations in this case are firms that do not pay dividends at all. Therefore, their exclusion is justified as they obviously do not try to attain any pay-out threshold. The majority of dividend pay-outs are concentrated in the first ten bins, i.e., up to 5% of previous year's total assets. Nevertheless, we keep the analysis of DIV in both versions as a reference. Similarly, in NSCF, it is practically never the case that the three components would sum up to exactly zero, meaning that zero observations are those of zero values in all three components and these again are validly excluded.[11] This is not as straightforward in scaled changes of dividends and net shareholder cash flows. D_DIV or D_DIV_DIV equal to zero may indicate a non-payer, but it can also indicate a no-change in dividends, keeping their level unchanged from the previous year. Analogously, D_NSCF and D_NSCF_NSCF values of zero can mean non-payers, no-changes in the sense that the firm only pays dividends and does not use repurchases and/or issuances or rare cases of the NSCF components summing exactly to zero.

We also separately evaluate bin(10) of the D_DIV_DIV distribution. The value of the test statistics of the GRPV test amounts to 6.22 and is significant at the 5% level. As the bin corresponds to a 10% to 11% increase of the dividends from the previous year, it also looks like a convenient orientation value for possible future pay-out increases. The GRPV

[11] Actually, there are seven cases in which dividends, repurchases and issuances sum up to exactly zero, but only pairwise – in none of them all three at the same time.

test value in bin(10) of the variable D_NSCF_NSCF is 0.33, limiting previous reasoning to cash dividend pay-outs only.

Focusing back on central bins, in Table 5 we investigate statistically significant (a 5% level is tested) differences between mean (median) values of discretionary accruals from the modified Jones model across bins. For each variable, with and without zeros, mean and median discretionary accruals from the model were computed for each bin in range from (-10) to (10), representing ±5% of lagged total assets or ±10% of lagged pay-outs, the difference due to different bin widths in the two approaches. Only the values for bins from (-2) to (2) are tabulated. We do this, firstly, because this is where our research interest lies as these are the most likely places in the distributions of pay-outs where the discretionary component of accruals would be important. Secondly, because there are not many significant differences further away from the centres of distributions. Finally, we keep our analyses compact for brevity of exposition. Bin means (medians) of discretionary accruals are compared to the means (medians) of discretionary accruals in the next bin using a t-test for the means and the Wilcoxson rank-sum test for the medians. For example, a boldfaced mean of DIV in bin(0) (0.955) indicates that it is significantly different from the mean in bin(1) (0.032). Similarly, a boldfaced median for NSCF in bin(-1) (0.011) indicates that it is significantly different from the median in the following bin(0) (0.091).

Note that seemingly missing values are actually excluded for clarity. Variable DIV does not have negative bin observations (no negative dividends), while the results for bins (-2), (1) and (2) are not listed for versions of variables without zero observations because they are exactly the same as on the left-hand version. The versions only differ in the number of observations in the central bin (bin(0)) and possible differences only arise in comparisons of bin(-1) to bin(0) and of bin(0) to bin(1).

In almost all instances significant differences in both means and medians of discretionary accruals are found at bin(0) or bin(-1) – the two that compare the central bin(0) with the neighbouring bins. Bin means of discretionary accruals are generally much larger than medians of discretionary accruals as a consequence of skewed distributions and are usually biggest in bin(0), means of bin(0) in first four variables being much bigger than means of other bins. Interestingly, excluding zero observations results in smaller bin(0) mean and median discretionary accruals compared to cases with all observations included and with the last two variables (D_DIV_DIV and D_NSCF_NSCF) they even become insignificantly different to other bins' means and medians. Assuming that discretionary accruals are associated with some form of purposeful managerial actions, and may be a tool to manage earnings or some other operating result by the management, their size and significance in central zero bins of distribution is at least indirect evidence of such actions.

Table 5: *Means and medians of discretionary accruals by bins*

Bin	Mean	Median	Mean	Median	Mean	Median	Mean	Median
	DIV		DIV (without 0)		NSCF		NSCF (without 0)	
-2					**0.169***	**0.056***		
-1					**0.114***	**0.011***	**0.114***	**0.011***
0	**0.955***	**0.128***	**0.006***	**-0.025***	**0.699***	**0.091***	0.011	**-0.026***
1	0.032	-0.012			0.027	-0.007		
2	0.038	-0.008			0.028	-0.010		
	D_DIV		D_DIV (without 0)		D_NSCF		D_NSCF (without 0)	
-2	**0.025***	-0.030			**0.063***	-0.018		
-1	**0.003***	**-0.035***	0.003	**-0.035***	**0.028***	**-0.020***	0.028	-0.020
0	**0.564***	**0.034***	**0.005***	**-0.023***	**0.347***	**0.004***	0.025	-0.022*
1	**0.049***	**0.006***			**0.035***	**-0.007***		
2	0.099	**0.028***			0.071	0.007		
	D_DIV_DIV		D_DIV_DIV (wo 0)		D_NSCF_NSCF		D_NSCF_NSCF (wo 0)	
-2	-0.012	-0.039			0.056	-0.028		
-1	**-0.022***	**-0.051***	-0.022	-0.051	0.099	**-0.015***	0.099	-0.015
0	**0.054***	**-0.016***	-0.023	-0.043	0.102	**0.007***	-0.002	-0.036
1	-0.007	-0.037			0.051	-0.038		
2	-0.007	-0.035			0.015	-0.033		

Notes: This table reports means and medians of discretionary accruals form the modified Jones model by central bins of distributions. Each variable has bin means reported in the first column and bin medians in the second column of its box and is firstly evaluated with all observations included and then with zero observations excluded ("wo 0" in the last variable row stands for "WITHOUT 0").

Bolded font and asterisk denote that respective means (medians) are different from means (medians) in the following bin at the 5% significance level, i.e., a bolded* mean (median) in bin(0) is different from the mean (median) in bin1 at 5%. Tests used were t-test for means and Wilcoxson rank-sum test for medians.

Modified Jones model is of the form: $NDACC = \alpha_1(1/TA_{t-1}) + \alpha_2(\Delta REV_t - \Delta REC_t) + \alpha_3(PPE_t)$. $NDACC$ are non-discretionary accruals, TA_{t-1} are lagged total assets (WC02999), ΔREV_t is the change in revenues (WC01001), scaled by lagged total assets, ΔREC_t is the change in receivables (WC02051), scaled by lagged total assets and PPE_t is gross property plant and equipment (WC02301), scaled by lagged total assets. To estimate α_1, α_2 and α_3 total accruals are considered as the dependent variable and calculated as: $TACC = (\Delta CA_t - \Delta CL_t - \Delta Cash_t + \Delta STD_t - Dep_t)TA_{t-1}$. ΔCA_t is the change in current assets (WC02201), ΔCL_t is the change in current liabilities (WC03101), $\Delta Cash_t$ is the change in cash and cash equivalents (WC02001), ΔSTD_t is the change in short term debt (WC03051), Dep_t are depreciation and amortization charges (WC01151) and TA_{t-1} are lagged total assets. Finally, discretionary accruals are obtained as the difference between total accruals and non-discretionary accruals predicted by the model.

DIV = dividends (WC04551) scaled by lagged total assets; NSCF = (dividends + repurchases (WC04751) – issuances (WC04251)) = net shareholder cash flows scaled by lagged total assets; D_DIV = change in dividends from year (t-1) to (t) scaled by lagged total assets; D_NSCF = change in net shareholder cash flows from year (t-1) to (t) scaled by lagged total assets; D_DIV_DIV = change in dividends from year (t-1) to (t) scaled by lagged dividends; D_NSCF_NSCF = change in net shareholder cash flows from year (t-1) to (t) scaled by lagged net shareholder cash flows.

For variables DIV, NSCF, D_DIV and D_NSCF bin width is 0.005 with lower bound inclusion, i.e., "zero bin" includes x if $0 \leq x < 0.005$, "bin one" includes x if $0.005 \leq x < 0.01$ etc., and for variables D_DIV_DIV and D_NSCF_NSCF bin width is 0.01 with lower bound inclusion, i.e., "zero bin" includes x if $0 \leq x < 0.01$, "bin one" includes x if $0.01 \leq x < 0.02$ etc. Bins in the range from -10 to 10 were tested but are not tabulated. Mean and median results of variables without zero observations are also not reported for bins -2, 1 and 2, as they are the same as in with zero observations included (the two versions differ only in the frequency of the zero bin).

The two signals combined, that of accruals and breaks in pay-out distributions, indicate that the thresholds identified in this study can be associated with some firms' management activity. As firms aim to meet their planned, announced or established levels of pay-out on one side, and face anticipations of shareholders and potential investors on the other side, thresholds in form of positive pay-outs or pay-out changes gain in importance. Not wanting to fail expectations firms may make use of accrual manipulation to arrive at desired financial results that enable a suitable pay-out policy.

5. ADDITIONAL TESTS

To address the potential sensitivity of discontinuity tests to neighbouring bin values suggested by previous research (Bennet & Bradbury, 2007), we first recalculate GRPV test statistics using two adjacent bins on either side of bin(0) (i.e., bins -2, -1, 1 and 2) and report it in column 1 of Table 6. The only difference to the main test is that the break in NSCF without zero observations is now only significant at 5% compared to previous 1% significance. All the other variables' τ values are very similar to previously reported ones. We also re-calculate the GRPV test using only next-to-adjacent bins (i.e., -2 and 2) and the results (not tabulated) remain quantitatively and qualitatively substantially unchanged. This confirms the robustness of earlier our results to the details of test specifications.

Extending the analysis beyond the primary hypotheses, we then use the test statistics to study what happens with the breaks in the distributions in relation to specified cut-offs, identified as potentially important for pay-out time dynamics. In columns 2 and 3 of Table 6 we look at the pre- and post- 2008 financial crisis periods. The inferences are unchanged with an adjustment in significance to 5% for NSCF and D_DIV_DIV, both without zero observations. What we do observe comparing the two sub-periods is that for the years 2008 and following all test values are smaller, mainly in the order of one half, than in pre-2008 period (apart from DIV and D_NSCF_NSCF, both without zero observations, which are insignificant as in the main test specification). Smaller values imply a less pronounced break in the distribution (although still highly significant) meaning less observations are concentrated in zero bins and more in the adjacent bins. This could be interpreted as some of the firms not pursuing or not being able to pursue pay-out thresholds in the crisis period, given the harsher economic conditions they found themselves in.

Table 6: *Additional GRPV discontinuity of distribution tests*

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	4 bins		crisis effect		ifrs used		finance act
	used	pre-2008	2008&→	no	yes	pre-1997	1997&→
DIV	349.19	297.96	251.13	274.89	278.36	74.53	400.21
DIV (wo 0)	2.28	0.82	1.61	1.03	1.07	2.97	0.23
NSCF	124.50	90.59	75.17	86.61	78.64	33.14	112.96
NSCF (wo 0)	8.39	9.88	5.62	8.94	7.10	0.19	12.75
D_DIV	401.72	281.52	201.50	261.58	227.93	109.76	335.98
D_DIV (wo 0)	155.96	114.15	47.71	107.18	62.14	68.41	103.21
D_NSCF	158.15	108.84	61.19	105.74	65.54	51.50	115.40
D_NSCF (wo 0)	86.81	58.07	23.54	56.50	27.28	33.03	53.32
D_DIV_DIV	78.00	69.00	24.20	66.44	30.39	52.07	52.27
D_DIV_DIV (wo 0)	23.69	20.58	5.56	19.44	8.42	18.54	12.57
D_NSCF_NSCF	39.52	36.85	15.56	36.82	15.64	27.51	29.31
D_NSCF_NSCF (wo 0)	1.75	1.41	1.77	1.46	1.47	2.83	0.31

Notes: Reported are τ values of GRPV discontinuity of distribution tests for zero bins of variables analysed with and without zero observations (the latter denoted by "wo 0" abbreviation). Column 1 reports statistics using 2 adjacent bins on either side of bin(0), columns 2 and 3 use 2008 as a cut-off year to analyse the effect of financial crisis, columns 4 and 5 analyse the effect of IFRS and column 6 and 7 use 1997 as a cut-off year to analyse the effect of change in legislation (Finance Act).
With H_0: *the distribution function is continuous*, the values of τ at standard levels of significance are: at 10% $|\tau| = 3.16$, at 5% $|\tau| = 4.47$, and at 1% $|\tau| = 10$. As the number of observations in adjacent bins is required by the test, in the first two rows (case of DIV) bins on the left of zero (negative bins) are empty (there are no negative dividends) and are as such affecting test statistic computation.
DIV = dividends (WC04551) scaled by lagged total assets (WC02999); NSCF = (dividends + repurchases (WC04751) – issuances (WC04251)) = net shareholder cash flows scaled by lagged total assets; D_DIV = change in dividends from year (t-1) to (t) scaled by lagged total assets; D_NSCF = change in net shareholder cash flows from year (t-1) to (t) scaled by lagged total assets; D_DIV_DIV = change in dividends from year (t-1) to (t) scaled by lagged dividends; D_NSCF_NSCF = change in net shareholder cash flows from year (t-1) to (t) scaled by lagged net shareholder cash flows.

Our next cut-off is IFRS implementation. International Financial Reporting Standards and their predecessors, International Accounting Standards, are mainly regarded as being of higher quality than existing local standards (e.g., Barth, Landsman, & Lang, 2008; Armstrong et al., 2010), although alternative views are also not uncommon (Soderstrom & Sun, 2007; Ahmed, Neel, & Wang, 2013), and they also directly affected accounting for dividends as noted under sample selection. IFRS are compulsory since 2005 and this appears as a ready candidate for assessing the standards' effects. We deem it a second-best option as before 2005 firms could voluntarily adopt IFRS and even after 2005 data shows some financial statements in our sample as being prepared under UK GAAP. Our database allows us to identify the standards which the company used in preparing its reports and we thus classify 7,678 observations as prepared under IFRS. These mainly coincide with the period after 2005, but there is some overlapping with local standards, especially in years 2004-2007. The results (columns 4 and 5 in Table 6) in terms of subsample comparisons are analogous to that for the crisis effect. IFRS observations exhibit notably lower test values than non-IFRS observations for all but two insignificant variables leading us to conjecture that IFRS usage is associated with "smoother" distributions. Potential explanation for this is the negative effect of stricter standards on firms' willingness and/or ability to achieve pay-out thresholds, positioning less of them in central bin(0).

We identify the last cut-off to be 1997 as pointed out by the dividend taxation literature. Namely, in order to end the discriminatory tax treatment in favour of dividend pay-outs compared to capital gains, the Finance Act of 1997 increased taxation of dividend income, primarily affecting pension funds that were the largest class of investors in UK equities.[12] Consequently, Bell & Jenkinson (2002) find a significant reduction in valuation of dividend income after the tax reform and initial evidence of reductions in dividend pay-out ratios, whereas Bond, Devereux & Kleem (2005) observe that it was the form of dividend payment that changed with the level marginally affected. Our two subsamples comparison in columns 6 and 7 of Table 6 reveals considerably smaller (yet again, still above critical values) values of discontinuity tests for most of the significant variables in the pre-1997 years compared to the later period. A potential explanation would be that after the 1997 tax reform dividends were not as large as before but still present (due to other investors' interests, signalling and other reasons), which resulted in their concentration in the smallest positive bin(s) of our distributions, producing a higher value of the test statistic. It has to be acknowledged however, that all these additional tests analyse only a specific factor possibly affecting pay-out dynamics and that firms' distributional decisions in real life are based on many elements, relative importance of which are changing in time. Moreover, even in our cases, there are overlapping effects especially towards the end of analysed time period.

6. CONCLUSION

This paper investigates the existence of pay-out-related thresholds as an extension of documented earnings management thresholds. With dividends and distinct net shareholder cash flows defined variables, discontinuities in their distributions are statistically analysed, employing a robust test that does not assume that the distributions of underlying variables are normal. The importance of pay-out policy for the firms' economic environment and for the firms' themselves (as a signalling mechanism, clientele and tax induced decisions etc.) leads us to consider threshold analysis to be of considerable importance for our study.

We find evidence of breaks in distributions at suggested thresholds of zero or zero change of variables in question, supporting our reasoning that these are important for firms. Dividend thresholds are more pronounced than net shareholder cash flows thresholds suggesting the dominating role of cash dividends over share buybacks and over the netting role of new shares' issuances. Although repurchases are almost as common as dividend pay-outs, their effect is much smaller. Adding share issuances in the calculation to

[12] More specifically, the 1997 Finance Act abolished repayment of dividend tax credits to tax-exempt investors, UK pension funds being the largest beneficiaries of the previous regulation.

arrive at net shareholder cash flows disperses the pay-out distributions and reduces the breaks. Hence, repurchases and issuances are relatively much less important drivers of targeted pay-out level in the broader sense and net shareholder cash flows do not represent a separate threshold independent of cash dividends.

Discretionary accruals as a proxy for earnings management are analysed at identified thresholds. We find significant differences and/or magnitudes of discretionary accruals at or in the closest proximity of central bins of distributions. This is another sign of their importance for firms as accruals are considered as a convenient and potentially strong earnings management tool. Additional analyses employ the discontinuity test to examine various sample partitions to arrive at more insightful results. We also find that a 10% dividend increase in the dividend paid is significant, suggesting the increase of dividends of 10% is common.

Known caveats relate to distributional analysis being questioned as an earnings management measure and, although supportive of our hypotheses and considered general, the accrual model employed is merely one of several accruals modes and these have been found to produce results of different significance or even conclusions. In a related, but not directly comparable research, Dechow, Richardson & Tuna (2003) are not able to confirm that discretionary accruals are driving the breaks in earnings distributions and offer supplementary explanations. Nevertheless, we consider the evidence in this paper strong enough to stress the importance of firms' pay-out policy, shedding additional light on the effects of pay-out policy components.

Finally, we also identify some potential areas for future research. For example, it might be possible to derive more precise tests that would be able to distinguish the effects of the financial crisis and the effects of new standards, where the two periods overlap significantly. This might be related to the use of more refined discretionary accruals models. These models might also be investigated independently of the breaks due to standards, financial crisis, etc. We also do not consider possible "real" earnings management (Roychowdhury, 2006), which might be a significant component of the overall management to achieve earnings and net shareholders' flows thresholds.

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