
Local and global illiquidity effects in the Balkans frontier markets

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We study market illiquidity across 11 national markets of the Balkans. In general, the EU member countries are more liquid than the nonmember countries. Turkey, however, has the most liquid market, while Serbia and Bosnia are the least liquid. Global illiquidity sourced from the US has a strong and positive impact on pricing in eight of the Balkans markets. In contrast, illiquidity transmitted from the EU impacts expected returns in only two instances, while local illiquidity is significant for just one market. Croatia and Slovenia are most susceptible to transmissions of regional illiquidity, each receiving illiquidity spillovers from four sources.

Keywords: market illiquidity; illiquidity transmissions; frontier markets; Balkans

JEL Classification: G15; G12; C59

I. Introduction

Next to the market risk premium found in the Capital Asset Pricing Model (CAPM), factors such as the firm size, book-market equity and momentum also take prominent places in the literature, see e.g., Fama and French (1992) and Carhart (1997). Illiquidity is still another factor, that is generally accepted to be important for pricing but which has largely been neglected in the literature. It is defined as a premium which discounts illiquid assets to prices lower than those of their liquid, but otherwise equivalent, counterparts (Amihud and Mendelson, 1986; Brennan and Subrahmanyam, 1996). Given that developed markets are relatively liquid a number of papers have turned to emerging markets to study illiquidity, which are expected to exhibit stronger illiquidity effects due to a lack of diversity in securities and ownership. In this article, we extend this argument and propose to analyse illiquidity in frontier markets that have not yet

reached the emerging market status. Since frontier markets provide diversification opportunities, it is important to understand which components of risk premia are priced according to their listed securities, see e.g., Berger *et al.* (2011) and Bekaert and Harvey (2002, 2003).

We investigate illiquidity across the following 11 markets of the Balkan Peninsula: Turkey, Greece, Bulgaria, Romania, Bosnia and Herzegovina (Sarajevo and Banja Luka), Croatia, Macedonia, Montenegro, Serbia and Slovenia. With the exception of Turkey and Greece, these countries form a homogenous group, having gone through similar stages of development since the Second World War. The last seven markets are of particular interest since they had come into existence following the breakup of Yugoslavia over the 1991 to 2006 period. As such, our *ex ante* expectation is to find relatively high illiquidity in these countries with a strong causal effect on the expected returns. On the other hand, the markets of Bulgaria, Greece, Slovenia, Romania and Croatia are

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members of the European Union and given the geopolitical risk characteristics of the region, we expected them to exhibit lower levels of illiquidity.¹ Our contribution to the existing literature is threefold – (i) we measure and compare illiquidity across the 11 markets of the Balkans, (ii) we present new evidence on the relevance of global (US and EU) and local illiquidity for asset pricing in these markets and (iii) we introduce the concept illiquidity spillovers, i.e., transmissions of illiquidity across markets, and test for its presence in the region.

As noted in Stoll (2000) and O'Hara (2003), the main issue present in the context of illiquidity measurement is the lack of relevant information, such as the bid-ask spreads. As a consequence, a number of alternative measures are developed in Roll (1984), Kyle (1985), Glosten and Milgrom (1985), Amihud and Mendelson (1986), Lesmond *et al.* (1999), Amihud (2002), Pastor and Stambaugh (2003) and Liu (2006), amongst others. Their effectiveness is investigated in Goyenko *et al.* (2009) who run a horse race of monthly and annual illiquidity measures. They conclude that ‘... the literature has generally not been mistaken in the assumption that liquidity proxies measure liquidity’ (p. 154).

Since the data availability for the markets under study is limited, we proxy illiquidity with the Zero Returns (ZR) measure introduced in Lesmond *et al.* (1999). This measure associates illiquidity with the proportion of days with ZR for each given month. There are two key justifications for the use of this proxy. First, zero-volume days, and hence ZR days, are more likely to be recorded for illiquid stocks. Second, higher transaction costs result in lower expected abnormal returns, and hence inhibit private information acquisition. Given that traders gather less private information on illiquid stocks, even when they trade on nonzero volume days, such stocks are more likely to record ZR days. Lesmond (2005) and Zhang (2010) find the ZR measure to be highly correlated with bid-ask spreads, and hence provide a reliable illiquidity proxy. This is corroborated in Goyenko *et al.* (2009), who present evidence suggesting the ZR proxy to be one of few measures whose performance does not deteriorate in the recent past.

Our study is formulated in a pricing model similar to that introduced in Amihud (2002). In order to evaluate the impact of domestic and global illiquidity, we control for the impact of a global equities factor. Thus, we generalize Amihud (2002) specification to an international setting. We also account for the possibility of measurement error in the ZR illiquidity proxy by introducing an errors-in-measurement model. As noted in Amihud *et al.* (2005), any constructed illiquidity proxy measures true illiquidity with error. The measurement error is expected to be significant in the case of the Balkans markets, which exhibit

significant frictions such as inadequate regulation, low trading volumes and frequency, unavailability and unreliability of information, as well as a short trading history. An additional statistical problem is the presence of unit-roots in some of the constructed ZR variables. We deal with this issue in two ways. First, in estimating the impact of illiquidity on expected returns, we difference the $I(1)$ ZR variables to induce stationarity. Second, when analysing illiquidity spillovers, since we wish all illiquidity variables to be in levels, we conduct our analysis in the Toda and Yamamoto (1995) framework. This allows us to test general parameter restrictions in vector autoregressions (VARs) which contain processes integrated to an arbitrary order.

We summarize our results as follows. We rank the markets from the most liquid to the least liquid in the following order: Turkey, Greece, Croatia, Slovenia, Bulgaria, Romania, Macedonia, Montenegro, Serbia, Bosnia (Sarajevo) and Bosnia (Banja Luka). With the exception of Turkey, which is a nonmember country, the EU countries rank relatively high in terms of liquidity confirming our *ex ante* expectations. While our findings regarding Croatian and Serbian markets are consistent with those of Benić and Franić (2008) and Minović (2012), the ordering of the remaining markets is new.

After controlling for the impact of a global equities factor and errors-in-measurement of illiquidity, we find a statistically significant impact of the US illiquidity to 8 of the 11 markets. The sign of the coefficients on the US illiquidity factor is positive in all cases, lending support to the hypothesis that illiquidity contributes to the risk premium (Amihud and Mendelson, 1986). In contrast, the EU and local illiquidity play minor roles in the pricing of the markets, with only one market (Romania) having a statistically significant local illiquidity effect. Montenegro appears to transmit illiquidity to four other countries, while Slovenia and Croatia receive foreign illiquidity in most instances. The direction of spillovers is similar to the findings of Hoti (2005), who examines return spillovers across the Balkans. Turkey, Greece, Romania, Bulgaria, Bosnia (Banja Luka) and Croatia exhibit a statistically significant relationship with a world equity factor, while other ex-Yugoslav countries do not. A lack of integration of frontier markets with the world market is discussed in Berger *et al.* (2011), Harvey (1995) and Bekaert and Harvey (1995).

The remainder of the article is set up in the following order. A brief literature review is presented in Section II, a first glance at the data in Section III, while the econometric method for assessing the impact of liquidity used in the subsequent analysis is outlined in Section IV. Section V contains empirical results and Section VI concludes.

¹ Croatia joined the EU in July 2013, while Macedonia, Montenegro, Serbia and Turkey are classified as candidate countries.

II. Literature Review

There is a growing body of empirical literature which lends support to the hypothesis that illiquidity positively impacts expected returns (Amihud and Mendelson, 1986). Given that comprehensive surveys of illiquidity measures are presented in Goyenko *et al.* (2009) and Amihud *et al.* (2005), we provide only a concise review and discuss a number of applied papers from the literature.

Measures of liquidity/illiquidity

Liquidity/illiquidity is unobserved and any study that investigates its impact must rely on a proxy constructed from available data. Two popular measures are the bid-ask spread and the turnover. While bid-ask spreads appear in the original papers by Glosten and Milgrom (1985) and Amihud and Mendelson (1986), amongst others, turnover is used in Datar *et al.* (1998), Nguyen *et al.* (2007), Rouwenhorst (1999), Claessens *et al.* (1995), Lesmond (2005), etc. Given that any constructed illiquidity proxy measures true illiquidity with error, see e.g., Amihud *et al.* (2005), new measures are continuously developed.

Liu (2006) provides a hybrid measure which adjusts turnover for the number of days when the daily trading volume is zero. His measure is said to account for additional dimensions of liquidity including trading volume, trading costs and the speed of trading. Similarly, Chai *et al.* (2010) define a measure, which is the summation of the standardized measures of three monthly trading characteristics: stock price, absolute monthly stock return and the thin trading measure proposed by Beedles *et al.* (1988). Amihud (2002) computes an illiquidity ratio which is defined as the ratio of absolute daily return to trading volume. It represents a measure of the price impact, since liquid securities can accommodate large trading volumes with small price concessions. It is applied in Acharya and Pedersen (2005) for the US market, Fang *et al.* (2006) for the Japanese market and Martínez *et al.* (2005) for the Spanish market, etc. The inverse of Amihud's (2002) measure is named Amivest, and is applied in Amihud *et al.* (1997) for testing of liquidity on the Tel-Aviv Stock Exchange.

In the case of emerging and frontier markets, access to information is limited and additional measures are computed from alternative variables. For instance, Lesmond *et al.* (1999) propose a measure of illiquidity, named ZR days, which is computed using only price series. ZR estimates illiquidity from the proportion of days with ZR in a given month. Price pressure (PP) of nontrading days used in Bekaert *et al.* (2007) is another similarly constructed measure. It represents a ZR measure with a price effect. Bekaert *et al.* (2007) note that lengthy periods of consecutive nontrading days should be associated with greater illiquidity effects than nonconsecutive periods.

However, the potential PP of any trade following a lengthy nontrading interval in a nonconsecutive period appears to result in a worse case of illiquidity. PP attempts to take this return 'catch-up' effect into account.

Empirical studies of liquidity/illiquidity

Brennan and Subrahmanyam (1996) contribute to the literature by investigating the relationship between illiquidity and monthly returns. They find a significant risk premium, which is associated with both the fixed and variable elements of the transaction cost. They further show that there is a concave relationship between the illiquidity premium and the variable cost, as originally postulated in Amihud and Mendelson (1986). Nevertheless, the relationship between the risk premium and the estimated fixed cost component is convex. In addition, Jacoby *et al.* (2000) provide evidence in support of the convex functional form.

Amihud (2002) broadens the illiquidity hypothesis and proposes that the expected illiquidity impacts expected returns over time as well as across companies, which was initially postulated. He finds evidence suggesting that expected returns are an increasing function of the expected illiquidity in the US. Acharya and Pedersen (2005) examine a simple equilibrium model with liquidity risk. In their liquidity-adjusted CAPM (LCAPM), a security's required return depends on its expected liquidity as well as on the covariances with the market return and liquidity. Acharya and Pedersen's model demonstrates that positive shocks to illiquidity, if persistent, are associated with low contemporaneous returns and high predicted future returns. The conditional version of the LCAPM is estimated by Acharya and Pedersen (2005) and Minovic and Zivkovic (2010).

Recently, the literature on liquidity effects has expanded to include studies of developing markets which exhibit higher levels of illiquidity. For instance, Claessens *et al.* (1995) examine the cross-sectional pattern of returns in emerging markets. They find that, in addition to market risk factor, liquidity proxies of size and trading volume also have significant explanatory power. Similarly, Rouwenhorst (1999) examines liquidity in 20 emerging markets using turnover as the proxy for liquidity. Hearn *et al.* (2010) apply an illiquidity measure proposed in Liu (2006) to study the African markets of South Africa, Kenya, Egypt and Morocco. They examine the impact of illiquidity and company size on asset pricing, in the context of the Fama-French factor model. Traditional factors as well as illiquidity are found statistically significant.

Currently there is only a limited body of literature on illiquidity in the Balkans region. As a part of a larger study, Lee (2011) shows that illiquidity risk is priced in Greece and Turkey. Bekaert *et al.* (2007) provide similar results

showing that liquidity has predictive power for future excess returns in these two countries, and that financial liberalization has not fully eliminated its impact. Benić and Franić (2008) measure and compare market illiquidity of Croatian equities to those of Bulgarian, Serbian, Slovenian, Hungarian, Polish and German markets over the period 2006 to 2008. They show that the Croatian market is more liquid than Bulgarian and Serbian markets, less liquid than Hungarian, Polish and German markets, and that it exhibits similar liquidity as the Slovenian market. Minović (2012) confirms that the Serbian market is less liquid than Croatian, although she demonstrates that both markets are highly illiquid.

III. Data

The data set consists of daily traded prices for common equities listed on the 11 Balkans markets obtained from Bloomberg, daily prices of US equities included in the Russell 3000 share market index,² daily prices included in the Bloomberg European 500 Index³ and monthly returns on a global equities index that is sourced from Ken French's online data library (http://mba.tuck.dartmouth.edu/pages/faculty/ken.french/data_library.html). In addition, monthly exchange rates are obtained from Datastream for the currencies involved. The exchange rates are used to convert the global equities factor, denominated in US dollars, to local currencies. The total number of shares listed on each of the national markets varies significantly and ranges from 139 firms listed in Slovenia to 1538 companies for Serbia. The US shares from the Russell 3000 index are used to compute the ZR measure of illiquidity, which we use as a proxy for global illiquidity. Finally, the sample interval spans from 3 June 2005 to 3 October 2012.

We construct daily log returns for each security j for the time period t as $r_{j,t} = 100 \times \log\left(\frac{P_{j,t}}{P_{j,t-1}}\right)$, where $\log(\cdot)$ is the natural logarithm, and $P_{j,t}$ represents the closing price of the security j . From these daily company returns, we construct monthly country returns by first forming daily equally weighted portfolio and then aggregating these country portfolios over time as follows $r_{i,t_M} = \frac{1}{N} \sum_{i=1}^N \sum_{t=1}^{T^*} r_{i,j,t}$. Here the index set $i = \{1, \dots, 13\}$ counts over the 11 Balkan markets, plus the US and the EU, $t_M = \{1, \dots, 12\}$ enumerates months, N represents the number of shares listed on each of the markets and T^* is the

number of trading days in a given month. We use equally weighted, rather than value-weighted indices due to the lack of data on company capitalizations. Even where market capitalizations are available they typically misrepresent the true firm size in frontier markets, as they are often based on stale prices of illiquid securities. Value-weighted indices are also dominated by large and relatively liquid firms, while there is only a small number of such companies in the frontier markets. Equally weighed country portfolios are also used in a number of seminal studies in the literature, see e.g., Amihud (2002), Pastor and Stambaugh (2003), Acharya and Pedersen (2005) and Sadka (2006).

A significant issue in estimating the effect of illiquidity on returns is the measurement of illiquidity, since it is unobserved. We follow Lesmond *et al.* (1999) to construct a ZR measure of illiquidity that uses only daily returns data. For each security j , we compute the ZR illiquidity proxy for month t_M as follows

$$\begin{aligned} ZR_{j,t_M} &= \frac{1}{T^*} \sum_{t=1}^{T^*} I(r_{j,t} = 0) \\ I(r_{j,t} = 0) &= \begin{cases} 1 & \text{if } r_{j,t} = 0 \\ 0 & \text{if } r_{j,t} \neq 0 \end{cases} \end{aligned} \quad (1)$$

Thus, illiquidity is associated with the proportion of trading days which have ZR. Once monthly ZR measures are constructed for each security, we create illiquidity proxies for each country i by taking the average across the securities listed on that country's exchange

$$ZR_{i,t_M} = \frac{1}{N} \sum_{j=1}^N ZR_{j,t_M} \quad (2)$$

When averaging in Equation 2, we censor those companies that trade less than 95% of the time.

Table 1 presents summary statistics for monthly market returns and illiquidity measures. The top performers over the period June 2005 to October 2012 are Bosnia (Sarajevo) with an annual average return of 29%, Romania with 26% and Turkey with 6.5%. Given that the time period covers the financial crisis period of 2007 to 2009, these markets have produced impressive yields. On the other hand, the worst performer is Bulgaria with an annual rate of return of -47.5%, followed by Macedonia (-36.3%) and Slovenia (-27.5%). Interestingly, the relationship between risk, as measured by SDs, and return does not follow the conventional 'high risk-high reward'

² This index measures the performance of the largest 3000 US companies, representing approximately 98% of the investable US equity market.

³ The Bloomberg European 500 Index is a free float capitalization-weighted index of the 500 most highly capitalized European companies.

Table 1. Descriptive statistics for monthly log returns and illiquidity measures

Country		Mean	SD	Skewness	Kurtosis	Jarque-Bera (<i>p</i> -value)	Autocorrelation $Q(5)$ (<i>p</i> -value)	Unit root (<i>p</i> -value)
Turkey	Return	0.065	0.299	-0.993	4.502	0.000	0.466	0.000
	ZR	0.167	0.034	-0.049	2.303	0.380	0.000	0.000
Greece	Return	-0.183	0.264	-0.588	4.923	0.000	0.062	0.000
	ZR	0.258	0.090	0.245	1.688	0.021	0.000	0.407
Romania	Return	0.258	0.398	1.583	9.264	0.000	0.000	0.000
	ZR	0.625	0.050	-0.135	3.678	0.352	0.000	0.004
Bulgaria	Return	-0.475	0.515	-0.354	3.312	0.310	0.000	0.000
	ZR	0.554	0.098	-0.285	2.333	0.221	0.000	0.395
Bosnia Sarajevo	Return	0.290	0.595	0.822	5.017	0.000	0.000	0.049
	ZR	0.693	0.075	-0.144	2.549	0.572	0.000	0.048
Bosnia Banja Luka	Return	0.030	0.562	0.044	7.410	0.000	0.000	0.037
	ZR	0.698	0.080	-0.275	2.606	0.407	0.000	0.136
Croatia	Return	0.026	0.251	1.244	7.479	0.000	0.000	0.000
	ZR	0.460	0.048	-0.266	2.731	0.498	0.009	0.000
Maced.	Return	-0.363	0.827	0.595	3.482	0.040	0.000	0.000
	ZR	0.666	0.083	-0.825	4.181	0.000	0.000	0.000
Monte.	Return	0.055	0.325	-0.269	10.593	0.000	0.000	0.000
	ZR	0.684	0.096	-0.384	2.965	0.314	0.000	0.100
Serbia	Return	-0.041	0.119	-0.773	13.494	0.000	0.387	0.000
	ZR	0.690	0.080	-1.442	14.738	0.000	0.439	0.000
Slovenia	Return	-0.275	0.276	-0.066	3.089	0.952	0.000	0.000
	ZR	0.480	0.054	0.439	2.610	0.164	0.000	0.212
EU	Return	0.061	0.201	-0.873	5.180	0.000	0.011	0.000
	ZR	0.028	0.024	1.027	3.907	0.000	0.003	0.016
US	Return	0.042	0.179	-0.768	4.706	0.000	0.049	0.000
	ZR	0.033	0.008	3.312	23.729	0.000	0.000	0.000

Note: Statistics are based on the period 3 June 2005 to 3 October 2012.

pattern. For instance, the worst performing countries of Macedonia and Bulgaria have about the same, or higher, levels of risk than Bosnia (Sarajevo) and Romania – the best performing markets.

All of the studied markets, with the exceptions of Bulgaria and Slovenia, show evidence of statistically significant non-normality at the 5% significance level according to the Jarque and Bera (1980) tests. This is mainly due to excess kurtosis in the distribution, as most of the skewness estimates lie close to zero. Most of the returns also exhibit serial correlation as indicated by the Ljung–Box statistics for the first five autocorrelations. The exceptions are Turkey and Serbia, in which cases we accept the null of no serial correlation as the 5% level. Lastly, according to the augmented Dickey and Fuller (1979) tests all 11 markets exhibit stationary returns series.

Turning to the statistics for the computed ZR illiquidity measures, we observe that they average from 0.17 for Turkey to 0.69 for Bosnia (Banja Luka). We may compare this to the average global illiquidity of 0.03. Dispersion of illiquidity is typically of smaller magnitude than that of returns, as is kurtosis. In fact Jarque–Bera tests suggest that 8 of the 11 markets have normally distributed ZR measures. The exceptions are Greece, Macedonia and Serbia. Interestingly, in five instances we find evidence of unit-root in the ZR variables at the

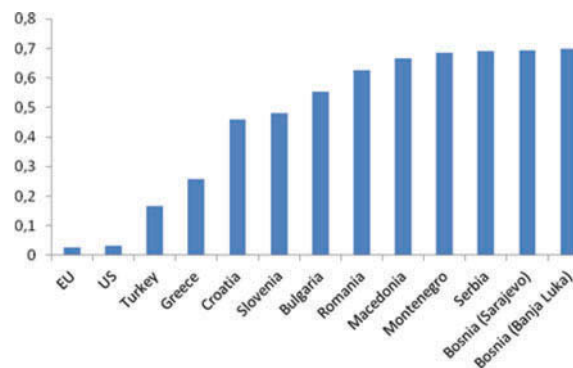


Fig. 1. Average illiquidity over the period 2005 to 2012

5% level. These are for Greece, Bulgaria, Bosnia (Banja Luka), Montenegro and Slovenia. The pattern of the average ZR measures presented in Fig. 1 shows that we may classify the 11 markets into 3 main groups according to their level of illiquidity. Turkey and Greece are clearly the most liquid markets, followed by Croatia, Slovenia and Bulgaria. The most illiquid markets comprise Romania, Macedonia, Montenegro, Serbia and Bosnia. The EU countries of Greece, Croatia, Slovenia, Bulgaria and Romania form a contiguous sequence of least illiquid markets.

Our conjecture is that information asymmetry contributes significantly to the levels of illiquidity in the Balkan markets. Akerlof (1970) provides arguments for the inverse relationship between information asymmetry and liquidity, and some empirical findings concerning this relationship in emerging markets are given in Zhang (2010). Although there is limited evidence on information asymmetry in the Balkan markets, Šoškić and Živković (2007) find that in the case of Serbia information asymmetry stems from the lack of regulatory control. In the absence of appropriate regulation governing information releases insider trading becomes widespread, and a large portion⁴ of the total capitalization turns illiquid.

IV. Econometric Methodology

We present our approach to measuring the impact of possibly mismeasured illiquidity first, and then discuss tests for illiquidity spillovers across the Balkans region.

Measuring the impact of illiquidity

Our model is formulated as an international factor pricing model which accounts for four factors: (i) world equity market factor, (ii) US illiquidity, (iii) EU illiquidity and (iv) domestic illiquidity. Consider the following pricing equation for the national equity returns

$$r_{i,t} = c_{0,i} + \beta_{0,i} r_t^{world} + \delta_{0,i} Illiq_t^{US} + \psi_{0,i} Illiq_t^{EU} + \gamma_{0,i} Illiq_{i,t} + \zeta_{i,t} \quad (3)$$

where $r_{i,t}$ represents return on the market of country i , r_t^{world} is the return on a world market index, $Illiq_t^{US}$ represents the US illiquidity, $Illiq_t^{EU}$ is the EU illiquidity and $Illiq_{i,t}$ accounts for local illiquidity in country i . The parameter vector $\lambda_0 = \{c_{0,i}, \beta_{0,i}, \delta_{0,i}, \psi_{0,i}, \gamma_{0,i}\}$ represents the true data generating process parameters. Our objective here is test for the impact of global and domestic liquidity, and to assess the level of regional integration with the world market. Unfortunately, $Illiq_{i,t}$ is unobservable and, as described in Sections I and II, $ZR_{i,t}$ measure is used as its proxy. Although $ZR_{i,t}$ has been shown to be an accurate representation of the true illiquidity (see e.g., Goyenko *et al.*, 2009), it is still a proxy and contains measurement error that needs to be taken into account. Further, its performance is likely be worsened in the face of market frictions encountered in the relatively undeveloped Balkan markets.

To allow for the possibility of measurement error, we assume that $ZR_{i,t} = Illiq_{i,t} + \eta_{i,t}$ with $E(Illiq_{i,t} \eta_{i,t}) = 0$

and $E(\zeta_{i,t} \eta_{i,t}) = 0$. Thus, $\eta_{i,t}$ represent the error in measurement of true illiquidity, which is assumed to be uncorrelated with illiquidity, as well as with the error in the pricing Equation 3. Should we in this instance attempt to estimate Equation 3 by a regression of the following form

$$r_{i,t} = c_i + \beta_i r_t^{world} + \delta_i ZR_t^{US} + \psi_i ZR_t^{EU} + \gamma_i ZR_{i,t} + \kappa_{i,t} \quad (4)$$

we would not be able to obtain an unbiased estimate of the parameter vector λ_0 . The reason is that the crucial assumption of $E(\kappa_{i,t} | r_t^{world}, ZR_t^{world}, ZR_{i,t}) = 0$ does not hold in Equation 4, see e.g., White (2001, p. 7). However since $E(\kappa_{i,t} | r_t^{world}, ZR_{t-1}^{world}, ZR_{i,t-1}) = 0$ under the assumption that $\eta_{i,t}$ is a zero mean white noise process, we may unbiasedly estimate

$$r_{i,t} = c_i^* + \beta_i^* r_t^{world} + \delta_i^* ZR_{t-1}^{US} + \psi_i^* ZR_{t-1}^{EU} + \gamma_i^* ZR_{i,t-1} + e_{i,t} \quad (5)$$

Provided that illiquidity is persistent, $ZR_{i,t-1}$ will be a good predictor for $ZR_{i,t}$, and will also act as an instrument for $Illiq_{i,t}$. According to the last column of Table 1, there are five instances when the Augmented Dickey–Fuller (ADF) test fails to reject the null of unit-root in the ZR measure. In those cases, we difference the series in order to make them stationary prior to estimating Equation 5.

Lastly, in order to account for possible autocorrelation and time-varying volatility, we augment Equation 5 by an autoregressive-moving average (ARMA) specification in the mean, and a time-varying variance as follows

$$\Phi(L)r_{i,t} = c_i^* + \beta_i^* r_t^{world} + \delta_i^* ZR_{t-1}^{US} + \psi_i^* ZR_{t-1}^{EU} + \gamma_i^* ZR_{i,t-1} + \Theta(L)e_{i,t} \quad (6)$$

In Equation 6, $\Phi(L) = 1 - \phi_1 L - \dots - \phi_p L^p$ represents the AR lag polynomial and $\Theta(L) = 1 - \theta_1 L - \dots - \theta_q L^q$ is the MA lag polynomial. The ARMA component of the model serves the purpose of filtering out any predictable pattern that may exist in the returns and may represent a time-varying alpha in the context of the CAPM. We select the correct ARMA order by minimizing the Aikake Information Criterion (AIC).

Given the extensive literature on time-varying volatilities in financial data we also model conditional variance. Formally, we specify the conditional distribution of the residuals as $e_{i,t} | \mathfrak{S}_{i,t-1} \sim N(0, h_{i,t})$, where the information set $\mathfrak{S}_{i,t-1}$ is a filtration generated by the innovations $e_{i,t-1}, e_{i,t-2}, \dots$, and $N(0, h_{i,t})$ denotes the normal distribution with zero mean and a time-varying variance $h_{i,t}$.

⁴It is estimated that in the Serbian market less than 10% of the total market capitalization is liquid (Minovic and Zivkovic, 2010; Minvic, 2012).

Thus, the unexpected return for each market in our data set is assumed to be conditionally normal, and to exhibit time-varying volatility. In order to model the conditional variances $h_{i,t}$ we use Bollerslev's (1986) generalized conditional heteroscedasticity (GARCH) model:

$$h_{i,t} = \omega_i + \alpha_i e_{i,t-1}^2 + \beta_i h_{i,t-1} \quad (7)$$

The parameters α_i and β_i capture the impacts on of previous period's unanticipated news shocks and volatility, respectively.

Regional illiquidity spillovers

In order to map out the channels of illiquidity transmissions across the region we conduct Granger causality tests. Let y_t be a 12×1 vector of the ZR illiquidity proxies for the 11 Balkans markets plus the global ZR variable. Granger (1969) motivates the concept of causality with the observation that an effect cannot precede a cause.⁵ Thus, if illiquidity in market A transmits illiquidity to market B , the former should help forecast the latter. In other words ZR_A Granger causes ZR_B if the mean squared of the forecast of ZR_B increases when we exclude ZR_A from the conditioning set.

Testing for Granger causality is relatively easy to implement in a VAR. We specify a p th-order VAR for the vector y_t as follows

$$y_t = c + \phi_1 y_{t-1} + \dots + \phi_p y_{t-p} + \varepsilon_t \quad (8)$$

In Equation 8, ϕ_1, \dots, ϕ_p are 13×13 coefficient matrices, c is a 13×1 vector of intercept terms and $\varepsilon_t = (\varepsilon_{1t}, \dots, \varepsilon_{13t})'$ is a zero mean white noise process. In this framework, noncasualty can be evaluated by testing the zero restrictions on the coefficient matrices ϕ_1, \dots, ϕ_p . For instance testing that y_t does not Granger cause y_k is done by conducting a Wald test of the following hypotheses for all market pairs (k, l)

$$\begin{aligned} H_0 : \phi_{j,kl} &= 0 \quad \text{for all } j = 1, \dots, p \\ H_1 : \phi_{j,kl} &\neq 0 \quad \text{for at least one } j \end{aligned} \quad (9)$$

A complication to the above analysis is introduced by our finding that some ZR series exhibit unit-roots, as illustrated in Table 1 and discussed in the section 'Measuring the impact of illiquidity'. The asymptotic Wald tests applied in Equation 9 are only valid if $y_t \sim I(0)$, i.e., y_t is a vector of variables integrated of order 0. Sims *et al.* (1990) and Toda and Phillips (1994)

show that the usual Wald statistics have a nonstandard asymptotic distribution if the process is $I(1)$. Such tests must be implemented through simulations of the limiting distributions under the null hypothesis. Thus, we may either difference the $I(1)$ variables in order to make them stationary as in the section 'Measuring the impact of illiquidity' or apply an alternative testing procedure that allows for a mix of $I(1)$ and $I(0)$ variables. For the purpose of testing for illiquidity spillovers, we prefer not to take the first approach as in that case y_t would comprise a mix of illiquidity levels as well as illiquidity differences, and the interpretation of results would become difficult. Therefore, we proceed to conduct our tests using the Toda–Yamamoto procedure.

Toda and Yamamoto (1995) suggest how to estimate VAR's specified in levels and test restrictions on parameters even if the processes contained in y_t are integrated or cointegrated of an arbitrary order. Their procedure is relatively simple in that it consists of finding the optimal lag length p using a usual selection procedure, e.g., (AIC), and then estimating a $(p + d_{\max})$ th-order VAR, where d_{\max} is the maximal order of integration that we suspect may occur in the process. Since in our case $d_{\max} = 1$ we estimate $(p + 1)$ th-order VARs. Granger causality tests are now conducted by testing the same hypotheses as in Equation 9, that is ignoring the $\phi_{(p+1)}$ th coefficient matrix. Following Rambaldi and Doran (1996), we implement this test in the Seemingly Unrelated Regression (SUR) framework.

V. Empirical Results

Table 2 presents the estimates of the pricing model described in Equation 6. As illustrated by the second and third columns of the table, illiquidity sourced from the EU, as well as local illiquidity, play only minor roles in the pricing of the Balkans markets. It appears that the EU illiquidity impacts pricing only in Bulgaria and Bosnia (Banja Luka), at the 5% significance. Moreover, in only 1 country of the 11 markets considered – Romania – is the local illiquidity a statistically significant factor. However, the sign of the estimated confidants in this case is negative, implying that higher illiquidity results in lower expected returns, a finding at odds with the theory and the evidence from developed markets. It is possible that in highly illiquid markets periods of excessive illiquidity prompt investors to unwind their positions and experience losses as they are unable to execute large transaction volumes

⁵ More formally, letting \mathfrak{S}_t be the information set consisting of all the relevant information available up to and including time t , $ZR_{A,t}$ Granger causes $ZR_{B,t}$ if $MSE(ZR_{A,t+h}|\mathfrak{S}_t) < MSE(ZR_{A,t+h}|\mathfrak{S}_t - ZR_{B,s} : s \leq t)$, where $MSE(\cdot)$ is the mean squared error of an h – step forecast of ZR_A conditional on the information set \mathfrak{S}_t . In the above specification, $\mathfrak{S}_t - ZR_{B,s} : s \leq t$ represents the information set \mathfrak{S}_t less the information contained in the past and present of the $ZR_{B,s}$ process.

Table 2. Measuring the impact of illiquidity

	US illiquidity δ_i^*	EU illiquidity ψ_i^*	Local illiquidity γ_i^*	Market integration β_i^*	ARMA (p, q) – GARCH (r, s)
Turkey	1.179 [0.023]	-0.289 [0.557]	0.350 [0.202]	0.596 [0.004]	(0,0)–(0,0)
Greece	2.937 [0.000]	0.193 [0.466]	-0.154 [0.423]	0.804 [0.000]	(0,0)–(0,0)
Romania	1.991 [0.008]	-0.024 [0.900]	-0.234 [0.005]	0.320 [0.000]	(2,2)–(1,0)
Bulgaria	8.142 [0.002]	-1.212 [0.029]	0.194 [0.228]	0.752 [0.001]	(0,1)–(0,0)
Bosnia (Sarajevo)	4.714 [0.108]	-0.244 [0.617]	-0.141 [0.487]	0.062 [0.772]	(4,3)–(0,0)
Bosnia (Banja Luka)	0.583 [0.015]	0.174 [0.000]	0.013 [0.587]	0.032 [0.041]	(3,2)–(1,0)
Croatia	2.681 [0.002]	-0.190 [0.456]	0.030 [0.800]	0.277 [0.000]	(1,0)–(0,0)
Macedonia	12.579 [0.001]	-0.992 [0.480]	-0.186 [0.566]	0.231 [0.582]	(4,3)–(0,0)
Montenegro	-1.942 [0.376]	-0.271 [0.331]	-0.007 [0.942]	0.064 [0.401]	(2,3)–(0,0)
Serbia	0.181 [0.240]	-0.049 [0.110]	0.015 [0.287]	0.002 [0.879]	(2,4)–(1,0)
Slovenia	3.867 [0.045]	-0.173 [0.685]	-0.110 [0.161]	0.216 [0.198]	(3,3)–(0,0)

Notes: The table reports the estimated coefficients of Equation 7, with the associated p -values computed with Newey–West HAC SEs given in brackets. The last column represents the orders of ARMA(p, q) and GARCH(r, s) components of the model. Shaded areas highlight coefficients statistically significant at the 5% level.

without affecting the price. This hypothesis is further discussed in Benić and Franić (2008).

In contrast to the EU and local illiquidity, the impact of the US illiquidity is significant and of correct, positive, sign in 8 of the 11 markets. Given that the coefficients on the local illiquidity and market integration are of similar magnitude, we may compare the magnitudes of coefficients on the US illiquidity across the markets as well.

Macedonia and Bulgaria appear to have the highest exposure to the US illiquidity with the estimated coefficients in excess of 12.6 and 8.1, respectively. The only three markets that do not appear to be impacted by the US illiquidity are Bosnia (Sarajevo), Montenegro and Serbia. Slovenia has a statistically significant coefficient of 3.9 to the US illiquidity. Greece, which has recently been at the centre of the sovereign debt crisis, has a statistically significant coefficient of 2.9, which when compared to the market beta of 0.80 illustrates the importance of illiquidity in the pricing of Greek equities. Our results are generally in line with the findings provided in Lee (2011), who demonstrates the importance of global illiquidity risk driven by the US. The only two markets that appear to be impacted by the EU illiquidity are Bulgaria with a negative coefficient and Bosnia (Banja Luka) with a positive coefficient. Nine of the markets do not appear to be impacted by the European ZR.

Six Balkan markets that receive a statistically significant impact from the global market factor are Turkey,

Greece, Romania, Bulgaria, Bosnia (Banja Luka) and Croatia. Interestingly, four of the five EU member states included in our study belong to this group. Of the non-EU countries, only Turkey and Bosnia (Banja Luka) exhibit a statistically significant relationship with the world factor. All of the six countries have betas less than 1, with Bosnia (Banja Luka) having the smallest beta of 0.03, while Greece and Bulgaria have the highest coefficients of 0.80 and 0.75, respectively. Interestingly, while we find the Greek market to be impacted by the global equities factor, Niarchos *et al.* (1999) report that it is not significantly related to the US market.

Our findings may shed light on some conflicting evidence regarding the integration of frontier markets with the world market. For instance while Berger *et al.* (2011) claim that frontier markets exhibit low levels of integration with the world market, Kenourgios and Samitas (2011) and Syriopoulos (2011) find the opposite. As evident from Table 2, Bosnia (Sarajevo), Macedonia, Montenegro, Serbia and Slovenia do not exhibit a statistically significant global market beta. Based on this criterion, these countries are not integrated with the global markets. However, this is only one piece of a multidimensional puzzle. Considering the coefficients on the US illiquidity we may conclude that all of Balkan markets, with the exception of Bosnia (Sarajevo), Montenegro and Serbia, are statistically integrated with the global markets via their dependence on global illiquidity. Thus it appears

that for these markets it is the global illiquidity premium, rather than other aspects of the global risk premium that investors price in the ex-Yugoslav countries.

The order of autocorrelation ranges across the markets from an ARMA(4,3) for Bosnia (Sarajevo) and Macedonia to ARMA(0,0) required for Turkey and Greece according to the AIC. The extent of time-varying volatility is limited in the Balkans' monthly returns. The first order autoregressive conditional heteroscedasticity model – GARCH(1,0) – fits well in three markets, while there is no evidence of GARCH effects in the remaining countries according to Engle (1982) LM GARCH test.⁶ This is a puzzling finding as one would expect these developing markets to have persistent GARCH effects over time.

Next we turn to discuss the results of illiquidity spillover tests presented in Table 3. Illiquidity spillovers are tested using the VAR model described in Equation 8, which is overfitted by one lag in order to deal with the problems introduced by some of the ZR variables having a unit-root (Toda and Yamamoto, 1995). According to the AIC a first order VAR is optimal, so when overfitted a VAR(2) is estimated. Spillover tests are conducted using the optimal lag length and ignoring the overfitted terms. In our application the optimal lag length is one, so we only test the elements of the first order VAR coefficient matrix.

Looking across the rows of Table 3, we see that Montenegro is a significant source of illiquidity in the region. There are three regional markets to which Montenegro transmits illiquidity with statistical significance, including Bosnia (Banja Luka), Serbia and Slovenia. Local illiquidity in Bulgaria, Macedonia and Montenegro is not Granger caused (at the 5% level) by illiquidity in any other country, although these countries Granger cause illiquidity elsewhere. Only Turkey does not transmit illiquidity to the region. On the other hand, local illiquidity in Turkey is Granger caused by illiquidity in Bosnia (Sarajevo) with a positive coefficient, and in Macedonia with a negative impact. Greece is a significant source of illiquidity for Croatia and Slovenia, with positive impacts. Romanian illiquidity has a negative statistically significant impact on the illiquidity in Croatia. Local Romanian illiquidity is Granger caused by European illiquidity with a negative coefficient. The effect of Bulgarian illiquidity is positive in Serbia while it is negative in the US market. Macedonian illiquidity has a negative statistically significant impact on the illiquidity in Turkey and positive statistically significant impact on the illiquidity in Bosnia (Sarajevo). Bosnian (Sarajevo) illiquidity has a positive statistically significant impact on the illiquidity in Turkey, and it has a negative statistically significant impact on the illiquidity in Slovenia. The effect of

Serbian illiquidity is negative in Bosnia (Banja Luka), while Serbian illiquidity is Granger caused by illiquidity in Bulgaria and Montenegro with positive coefficients. Local illiquidity in Croatia is Granger caused by illiquidity in Greece with a positive impact, and Romania and Slovenia with negative effects. Slovenian illiquidity Granger causes illiquidity in Croatia, while local illiquidity in Slovenia is Granger caused by illiquidity in Greece, Montenegro and the EU with positive impacts, and in Croatia with a negative effect.

The impact of European illiquidity is statistically significant only in two cases. It Granger causes illiquidity in Romania with a negative coefficient and in Slovenia with a positive coefficient. The impact of world illiquidity is also statistically significant in two instances. World illiquidity Granger causes illiquidity in Greece and Croatia, but with a negative coefficient in Greece and a positive parameter in Croatia. Another observation worth noting is that in 9 out of 11 instances, the diagonal elements of the table are statistically significant indicating persistence of local illiquidity. Only in the cases of Romania, Croatia, Slovenia and EU the autoregressive terms are not statistically significant.

Although the illiquidity transmission results presented here are new, and have not been reported before; they support the findings reported in Hoti (2005), who studies return spillovers across Albania, Bulgaria, Greece, Romania, Serbia, Montenegro⁷ and Turkey. Hoti concludes that return spillover patterns she uncovers indicate close ties between the Balkan countries in terms of their political, economic and financial affairs.

VI. Conclusion

We measure and study illiquidity across 11 equity markets of the Balkans Peninsula: Turkey, Greece, Bulgaria, Romania, Bosnia and Herzegovina (Sarajevo and Banja Luka), Croatia, Macedonia, Montenegro, Serbia and Slovenia. We compute ZR measure of Lesmond *et al.* (1999) as a proxy of illiquidity, given that true illiquidity is unobserved. It appears that the EU member countries of Greece, Slovenia, Croatia, Bulgaria and Romania have lower levels of illiquidity than most of the nonmember countries. However, Turkey – a nonmember country – has the most liquid market. These results are consistent with the findings reported in Syriopoulos (2011) and Kenourgios and Samitas (2011).

Next we formulate a pricing model that takes into account the possibility of measurement error in the constructed ZR variable. It appears that only 6 of the 11 markets studied here exhibit a statistically significant

⁶ Results of the GARCH test are available upon request.

⁷ In 2005 Serbia and Montenegro formed one country.

Table 3. Illiquidity spillover tests

		Granger causality to									
Granger causality from	Turkey	Greece	Romania	Bulgaria	Bosnia (Sarajevo)	Bosnia (Banja Luka)	Croatia	Macedonia	Montenegro	Serbia	Slovenia
Turkey	0.472 [0.000]	-0.204 [0.091]	-0.205 [0.249]	0.086 [0.629]	0.210 [0.418]	0.184 [0.352]	-0.024 [0.884]	0.379 [0.189]	0.079 [0.802]	0.381 [0.162]	-0.225 [0.179]
Greece	0.189 [0.082]	0.834 [0.000]	0.311 [0.066]	0.195 [0.246]	0.183 [0.455]	0.352 [0.060]	0.458 [0.003]	0.337 [0.217]	0.141 [0.638]	-0.043 [0.866]	0.387 [0.015]
Romania	0.024 [0.769]	0.109 [0.211]	0.104 [0.422]	-0.089 [0.487]	-0.132 [0.482]	-0.168 [0.241]	-0.292 [0.012]	-0.034 [0.871]	-0.091 [0.694]	-0.365 [0.065]	-0.076 [0.532]
Bulgaria	-0.007 [0.923]	0.054 [0.479]	0.128 [0.256]	0.491 [0.000]	0.011 [0.978]	0.189 [0.131]	0.059 [0.561]	0.019 [0.918]	0.008 [0.968]	0.285 [0.047]	-0.015 [0.890]
Bosnia (Sarajevo)	0.106 [0.039]	0.065 [0.234]	-0.013 [0.870]	0.035 [0.659]	0.304 [0.009]	-0.146 [0.101]	0.009 [0.904]	-0.103 [0.427]	-0.230 [0.059]	0.029 [0.811]	-0.190 [0.012]
Bosnia (Banja Luka)	0.009 [0.905]	-0.077 [0.326]	0.145 [0.319]	0.082 [0.475]	-0.245 [0.143]	0.368 [0.004]	-0.018 [0.864]	0.036 [0.845]	0.119 [0.563]	0.079 [0.651]	-0.080 [0.460]
Croatia	0.103 [0.162]	0.163 [0.038]	0.098 [0.399]	0.164 [0.156]	-0.083 [0.620]	0.137 [0.285]	-0.079 [0.447]	0.124 [0.501]	0.021 [0.928]	0.020 [0.908]	-0.100 [0.358]
Macedonia	-0.125 [0.011]	-0.083 [0.114]	0.013 [0.867]	-0.029 [0.707]	0.335 [0.003]	0.070 [0.415]	0.015 [0.826]	0.277 [0.028]	0.183 [0.187]	0.058 [0.623]	-0.078 [0.285]
Montenegro	-0.033 [0.428]	-0.014 [0.752]	-0.128 [0.053]	-0.122 [0.063]	-0.026 [0.789]	0.144 [0.048]	-0.000 [0.994]	0.040 [0.705]	0.372 [0.001]	0.194 [0.049]	0.129 [0.035]
Serbia	0.016 [0.697]	-0.034 [0.443]	0.048 [0.463]	-0.073 [0.262]	0.043 [0.654]	-0.163 [0.025]	-0.027 [0.640]	-0.121 [0.253]	-0.124 [0.288]	0.332 [0.000]	-0.020 [0.746]
Slovenia	-0.105 [0.161]	-0.007 [0.929]	-0.106 [0.366]	0.143 [0.222]	-0.100 [0.568]	-0.099 [0.446]	-0.265 [0.011]	0.097 [0.608]	-0.235 [0.258]	0.086 [0.632]	-0.014 [0.899]
EU	-0.079 [0.545]	-0.219 [0.112]	-0.424 [0.038]	-0.034 [0.868]	0.031 [0.916]	0.046 [0.839]	0.131 [0.473]	0.216 [0.511]	-0.153 [0.673]	-0.033 [0.914]	0.374 [0.050]
US	0.407 [0.638]	-1.989 [0.029]	-0.093 [0.945]	-1.245 [0.354]	-1.118 [0.568]	-0.829 [0.579]	3.183 [0.008]	1.240 [0.600]	-2.423 [0.312]	-1.712 [0.405]	0.539 [0.670]

Notes: The table presents the estimates of ϕ_1 given in Equation 8 estimated as a SUR, and their associated p -values. Shaded areas highlight coefficients statistically significant at the 5% level.

beta with respect to a world market index. Based on this finding, one may be tempted to conclude that about a half of the Balkan markets are not integrated with international equities. However, there is a strong and statistically significant impact of the US illiquidity on eight of the Balkan markets, which suggested that most of the region is in fact indirectly integrated with the US market via illiquidity spillovers. The sign of the estimated coefficients associated with the US illiquidity is positive in all cases, signifying that the US illiquidity is a priced risk factor, as suggested in Amihud and Mendelson (1986). In contrast, illiquidity sourced from the EU, as well as local illiquidity, is insignificant in most cases.

Investigating illiquidity spillovers across the region we find statistically significant transmissions across the region. Greece, Bulgaria, Bosnia (Sarajevo), Montenegro and Macedonia appear to be major transmitters of illiquidity. In terms of receiving illiquidity, the most susceptible markets are Slovenia and Croatia, which are affected by illiquidity from three other countries of the region, as well as from the global markets. Serbia, Bosnia (Banja Luka), Turkey and Greece also receive illiquidity from two other countries of the region. Interestingly, European and global illiquidity Granger cause illiquidity in only 2 of the 11 markets. Our results extend the findings reported in Hoti (2005) who investigates return transmissions in the Balkan markets.

Lastly, we provide a brief discussion regarding the prospects of Balkan equity markets in the context of illiquidity pricing. As demonstrated in Table 2, global (US) illiquidity appears to be the most significant factor in the pricing of the Balkans markets. This is likely due to the fact that the average level of illiquidity across the Balkan markets is high, which magnifies their susceptibility to any marginal change in global illiquidity. As these markets mature and integrate with international trends, we foresee two possible developments relating to the liquidity component of risk premium.

We expect the overall level of illiquidity in the Balkan countries to decline as more international investors enter the markets. Currently, a major obstacle to this happening is poor regulation and a lack of reliable information. Particularly concerning are information asymmetries caused by inadequate regulation that result in widespread insider trading (Šoškić and Živković, 2007). Nevertheless, as described in a recent report by Deloitte Private Equity (2013), steps are being made towards improving market conditions across the region. Assuming that the overall level of illiquidity eventually declines, the sensitivity of the Balkans countries to global illiquidity is expected to decrease, subsequently leading to lower risk premia. Second, counteracting this effect is the impact of increased market integration that may amplify transmissions of illiquidity from other international bourses, especially from the EU that currently exerts a negligible impact. The net

result of future improvements in market regulation, and increased levels of international integration is difficult to foresee. Nevertheless, changes in market structure and regulation take time, and in the interim the Balkan frontier markets are likely to continue to experience illiquidity as a major source of risk.

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