Compensation vouchers and equity markets: Evidence from Hungary

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Abstract

One of Hungary’s policies during the transition from a centrally planned to a market economy was the issue of compensation vouchers – a unique security designed both as a privatization mechanism and as a form of restitution for Hungarian citizens who suffered property losses in post-war nationalizations. The coupons were actively traded on the Budapest Stock Exchange (BSE). This paper examines the intertemporal behavior of the Hungarian voucher and equity markets in an effort to assess the efficiency of these markets and to gauge the degree of interaction between the two different assets. Evidence from variance ratio tests indicates that stock and voucher trading are each individually weakly efficient. Furthermore, vector autoregressions and cointegration methods show that there is little detectable intermarket interaction: a result which is consistent with joint efficiency. Thus, although the Hungarian equity market is small, it appears to function remarkably well. © 2000 Elsevier Science B.V. All rights reserved.

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1. Introduction

Among the many reforms implemented in transitional economies, the slowest changes have occurred in the area of financial market development. The banking sectors in most of these countries are still dominated by state-owned institutions while equity and bond markets typically remain small and illiquid. On the other hand, dramatic progress has been made in the privatization of state-owned enterprises and property, frequently through programs that included some form of freely distributed vouchers which were designed to facilitate the transfer of state assets to private ownership. In some countries such as Hungary and Russia, these vouchers were legally tradable in official security markets. The novelty of these markets raises the obvious question of whether voucher trading was efficient and whether such trading had an impact on nascent equity markets. The contribution of this paper is to shed light on these issues using data from the Budapest Stock Exchange (BSE).

An analysis of the Hungarian voucher market is important for the design and formulation of policies in transitional economies that may be contemplating voucher privatization programs. For instance a fundamental question that has faced all of the economies using vouchers as part of a privatization scheme is the problem of coupon tradability. In particular, one would like to know whether unfettered voucher trade has occurred efficiently or whether informational asymmetries are so widespread in these countries that trading is perverse. Moreover, it is clear that the development of equity markets has been affected by choices regarding privatization methods (Blommestein, 1998; Nivet, 1997). Therefore, it is useful to use the Hungarian experience to examine potential interactions between voucher and equity trading.

The fundamental question is: What are vouchers? On one hand, one could argue that they represent a (risky) claim to productive assets in the future. Viewed in this way vouchers might be considered as a type of equity or government bond.¹ For example, after its aborted “second wave” voucher privatization attempt in 1995, the Slovak government explicitly converted outstanding vouchers to five-year bonds guaranteed by the Slovakian National Property Fund (NPF) (see EBRD, 1996). Since these bonds can be sold or used to buy shares of companies held by the NPF, they are functionally equivalent to tradable vouchers. On the other hand, it is possible that the public might perceive vouchers to be entirely divorced from fundamentals other than the

¹ There is a superficial similarity between vouchers and convertible bonds since vouchers can be converted into equity. However, there are some important distinctions. In the case of a standard convertible bond, the conversion option is held by the bondholder while in the case of a privatization voucher, the government has a significant amount of control over the timing of the conversion. Furthermore, the actual conversion rate is known for a convertible bond whereas the voucher-to-share exchange rate is primarily determined by the privatization auction price.
behavior and commitment of the government. In this case, one would expect that voucher markets might operate in a quite different manner. For instance, early in the privatization process, there were concerns that the introduction of freely tradable vouchers would be seen by recipients as a (liquid) windfall of government transfers and would therefore exacerbate inflation via wealth effects.  

The remainder of the paper is organized as follows. Section 2 briefly discusses the development of the Hungarian equity and voucher markets. Section 3 provides a summary of the time-series methodology employed in this study and reviews previous applications of these techniques. Section 4 describes the data, followed by a presentation of results in Section 5. Finally, Section 6 concludes with a discussion and some closing remarks.

2. The Hungarian stock and voucher market

An important characteristic of the BSE and Hungarian policy in general is the fact that Hungarian privatization law permitted unfettered trading of vouchers, unlike many other transitional countries where possession and use of vouchers were restricted to the individual who initially received them. Other countries that have allowed legal trading of coupons include Estonia, Latvia, and Russia. In Russia’s case, the Russian Exchange, became an important center for the trading of privatization securities.  

Anecdotal evidence suggests that the trading of privatization securities was beneficial to the development of equity markets in both Hungary and Russia through the increased participation and experience of traders who handled the vouchers. For instance, the National Bank of Hungary (1995) states, “several hundred thousand citizens who have been given compensation vouchers have virtually been forced to pay attention to the Stock Exchange where the value of their vouchers is being determined.” In 1994, secondary market voucher transactions accounted for about 9% of all trading volume on the BSE.  

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2 See Lavigne (1995, p. 167) and Major (1993, p. 89). An interesting policy during the French Revolution provides a historical precedent to the current privatization and voucher episode. The French government nationalized church properties and issued notes (assignats) which were exchanged for outstanding government debt. While the assignats were intended to be tendered ultimately in auctions for church property and then immediately destroyed, they began to circulate as money. See Sargent and Velde (1995).

3 While the Russian coupon market would also make an interesting case study, the reliability and continuity of basic indicators such as stock price indices are highly suspect. See The Economist, 1994, March 12, p. 91.

4 Total trading volume on the BSE in 1994 – including all transactions of equities, government and corporate bonds, treasury bills, and compensation coupons – was Ft 105.5 billion while trading in compensation coupons was Ft 9.35 billion. National Bank of Hungary (1996, p. 126).
same year, 75% of trading volume on the Russian Exchange involved privatization vouchers. According to the Russian Exchange (1994) Annual Report, “the experience gained while dealing in (privatization) cheques was very helpful in creating the main mechanisms of the securities market.”

The BSE resumed operations in 1990, making it one of the first Eastern European bourses to “reopen”. Thus, the BSE has available a relatively long span of data required for the time-series techniques described in the next section. Partly due to its longevity, the BSE index – which is modeled after the S & P 500 and German DAX – is reported by most financial information systems. Moreover, trading on the BSE started two years before the initiation of the voucher program, so accounting and reporting mechanisms were already well established by the time coupon trading started. 5

Another attractive characteristic of the BSE is the large presence of foreign ownership, not only through direct institutional investment but also due to the unusual feature that shares in several Hungarian firms are simultaneously listed on the Vienna, Stuttgart, Munich, and London stock exchanges. It was estimated that in 1996, approximately 70% of the publicly traded stocks (the free float) on the BSE was owned by foreign entities such as large institutional investors. The World Bank (1996) argues that foreign demand in emerging security markets results in “infrastructure improvements” since institutional investors require good custody, trustee, audit, and bank payment systems. All of these reasons reinforce the impression that the quality of data from the Hungarian exchange is relatively high.

Despite the fact that it was one of the first stock exchanges to re-open in Eastern Europe, the Hungarian stock exchange remained small compared to some of the other markets, particularly in terms of the number of listed firms. The number of firms listed on the Czech and Slovakian exchanges literally exploded due to their large mass-privatization “waves” in the mid-1990s (see Table 1). On the other hand, the pattern of firm listings on the Polish stock exchange was closer to Hungary with a gradual increase in the number of companies. Table 1 also reveals that the BSE lagged the other three regional exchanges in terms of turnover ratios, and trading values in Hungary were generally lower than the Czech Republic and Poland but higher than Slovakia.

It is important to distinguish between “compensation” and “privatization” vouchers. Compensation vouchers were issued by the Hungarian government as restitution for losses suffered by citizens during nationalizations of private

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5 This stands in sharp contrast to other regional stock markets such as Prague and Bratislava which were created practically overnight as firms were converted to joint-stock corporations in preparation for their mass privatization programs.

6 National Bank of Hungary (1997b, p. 37). Creditanstalt (1994) reported that this percentage was as high as 85% in 1994.
property after the Second World War. Privatization vouchers, on the other hand, were securities issued as part of mass-privatization programs such as those in the Czech Republic, Slovakia, and Russia. As such, the securities were widely distributed to the population.\textsuperscript{7}

In Hungary, only those who could document that their property had been illegally taken from them by the state were eligible for compensation. In all, approximately 1.2 million Hungarians (10% of the population) received vouchers with a total face value of 120 billion Forints (US$1.1 billion).\textsuperscript{8}

Although exact data on the quantity and timing of voucher distribution is not publicly available, voucher issues started at the end of 1991 and continued throughout 1992. In December 1992, the BSE introduced the formal secondary market for vouchers.

By design, Hungarian compensation vouchers functioned in many interesting ways. In addition to being sold on the BSE or tendered in privatization

\begin{table}

\caption{Comparative annual data for Central European stock exchanges, 1991–97\textsuperscript{a}}
\begin{tabular}{lcccccccc}
\hline
\hline
\textbf{Czech Republic} & & & & & & & & \\
Capitalization/GDP (%) & – & – & – & 16.5 & 33.2 & 32.9 & 24.6 \\
Turnover ratio (%) & – & – & – & 22.4 & 23.2 & 46.6 & 55.2 \\
Trading value (US$ mn) & – & – & – & 1328 & 3630 & 8431 & 7055 \\
Number of firms & – & – & – & 1024 & 1635 & 1588 & 276 \\
\textbf{Hungary} & & & & & & & & \\
Capitalization/GDP (%) & 1.5 & 1.5 & 2.1 & 3.9 & 5.4 & 11.8 & 32.8 \\
Turnover ratio (%) & 23.2 & 6.8 & 12.2 & 16.8 & 14.8 & 31.1 & 51.3 \\
Trading value (US$ mn) & 117 & 38 & 99 & 270 & 355 & 1641 & 7684 \\
Number of firms & 21 & 23 & 28 & 40 & 42 & 45 & 49 \\
\textbf{Poland} & & & & & & & & \\
Capitalization/GDP (%) & 0.2 & 0.3 & 3.1 & 3.3 & 3.8 & 6.2 & 9.0 \\
Turnover ratio (%) & 19.4 & 75.2 & 80.2 & 167.9 & 60.7 & 66.0 & 65.7 \\
Trading value (US$ mn) & 28 & 167 & 2170 & 5134 & 2770 & 5538 & 7977 \\
Number of firms & 9 & 16 & 22 & 44 & 65 & 83 & 143 \\
\textbf{Slovakia} & & & & & & & & \\
Capitalization/GDP (%) & – & – & – & 7.9 & 7.1 & 11.5 & 9.4 \\
Turnover ratio (%) & – & – & – & 11.0 & 67.4 & 106.4 & 118.6 \\
Trading value (US$ mn) & – & – & – & 120 & 832 & 2321 & 2165 \\
Number of firms & – & – & – & 18 & 18 & 816 & 872 \\
\hline
\end{tabular}

\textsuperscript{a}Turnover ratio is the total value traded divided by annual market capitalization. \\

\textsuperscript{7} Securities resembling compensation or privatization coupons have also been issued in Bulgaria, Estonia, Latvia, Poland, and Romania.

\textsuperscript{8} Creditanstalt (1994). Comparable figures are reported by EBRD (1994).
auctions, they could also be used in the purchase of state-owned land and dwellings, or could be exchanged by the elderly for a life insurance annuity which was expected to be funded by the State Property Agency (SPA) through privatization income. Moreover, there was evidence of a growing informal market for coupons: shops began to accept the securities at a discount in exchange for goods, and advertisements for buying and selling vouchers appeared in local newspapers (Fletcher, 1995).

The trading history of these compensation vouchers is summarized in Table 2. Relative to the value of equities traded, there was initially strong interest in the voucher market followed by a gradual decline. This pattern is also detected in the turnover ratios for the two assets. While the turnover ratio for vouchers remained below 20%, the same indicator for equities by 1997 climbed to 51% (see Table 1).

It should be emphasized that in many cases such as the Czech, Slovak, and Russian examples, distributions of privatization vouchers in support of mass-privatization programs were much more extensive than the Hungarian compensation scheme. In the former cases, 80–90% of the eligible population received some form of the security. Nevertheless, in practice both compensation and privatization coupons functioned as a method of transferring state assets to private hands. In Hungary, the State Privatization Agency typically reserved a tranche of shares in a privatized firm for voucher purchases. Moreover, it was entirely expected that the vouchers would first be transferred to entrepreneurs in secondary markets and subsequently used for the purchase of state assets (Koranyi, 1993).

Since there is no previous literature on the relationship between voucher and equity markets, it is difficult to anticipate the type of interactions that might exist between these markets. However, if we view vouchers as a type of government bond, the large body of literature analyzing the co-movement of bond and equity prices is relevant. The theme in this research is that returns on the two assets react to common outside influences such as changing discount rates (Shiller, 1982; Shiller and Beltratti, 1992), cyclical business conditions (Fama and French, 1989), or changes in risk premia (Barsky, 1989). In a related study examining the relationship between bond and equity markets, Fama and

<table>
<thead>
<tr>
<th>Table 2</th>
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<tbody>
<tr>
<td>Hungarian equity and compensation voucher trading, 1991–97</td>
</tr>
<tr>
<td>-----------------</td>
</tr>
<tr>
<td>Value of shares traded</td>
</tr>
<tr>
<td>Value of vouchers traded</td>
</tr>
<tr>
<td>Voucher/Share ratio (%)</td>
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<tr>
<td>Voucher turnover ratio (%)</td>
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</tbody>
</table>

French (1993) find that stock market returns are linked to bond market factors and they show that the relationship is one-way – factors important for explaining bond returns capture common variation in stock returns but stock-market factors play little role in returns on government and corporate bonds. This result is also supported by Rahman and Mustafa (1997) who find Granger causality from bond markets to equity markets.

In transitional economies, it is likely that government policy was an important exogenous influence on asset prices. The progress of privatization programs throughout Central Europe was considered an important signal that governments were committed to market-oriented reforms. In Hungary this was particularly true because privatization auctions in Hungary allowed for bids in vouchers and cash. Thus, revenues generated by the sale of state-owned enterprises were considered an important source of financing for both the government deficit and the external foreign debt, and the usefulness of the vouchers depended on the future supply of state assets through the privatization program. Furthermore, it is widely believed that the overall efficiency of public and private enterprises improves as hardening budget constraints and increased competition force better allocation of resources throughout the economy. For all these reasons, it seems plausible that this information would affect the behavior of both the voucher and equity market.

3. Background and methodology

The “efficient market hypothesis” is attributed to Samuelson (1965) and Fama (1970) who categorized market efficiency as weak, semi-strong, or strong, depending on the hypothesized information set available to agents. 9 Fama (1991) offered a separate (three-category) taxonomy of efficient-market tests consisting of (1) tests for return predictability, (2) event studies, and (3) tests for private information. This paper employs several time-series techniques which can be classified in the first category: tests for return predictability. The underlying premise motivating techniques in this category is that if security prices reflect all available information, then market participants should be unable to form meaningful forecasts of future returns. Assuming that agents are limited to using only the information in past prices, then return unpredictability corresponds to Fama (1970) version of weak-form efficiency.

9 It is not my intention to contribute to the debate concerning the definition of an efficient market since this would take me away from the empirical study at hand. I ascribe to Fama (1991) view that meaningful research “improves our ability to describe the time-series and cross-section behavior of security returns.” Readers interested in the evolution of the efficient market hypothesis are referred to LeRoy (1989) and Fama (1991).
In a univariate context, if a price sequence, \( \{x_t\} \), follows a random walk, we can conclude that the evidence is consistent with the joint hypothesis of unpredictable returns and constant expected equilibrium returns. I will maintain this joint hypothesis as my working definition of weak efficiency. Univariate tests conducted here include unit root tests due to Dickey and Fuller (1979) and Zivot and Andrews (1992), Vogelsang’s (1997) structural change test, and Lo and MacKinlay’s (1988, 1989) version of the variance ratio test for random walks.

Variance ratio tests have been used extensively in the financial literature. Lo and MacKinlay (1988) and Poterba and Summers (1988) are representative studies that use the technique to examine market indices and firm-level price data from the NYSE. Lo and MacKinlay reject the random walk hypothesis for weekly stock market returns while Poterba and Summers provide evidence that monthly NYSE returns exhibit positive autocorrelations over one month horizons and negative correlations (mean reversion) over longer, yearly intervals. However, they are unable to statistically reject the random walk hypothesis at conventional levels. Examples using data from other financial markets include Lee (1992), Ayadi and Pyun (1994), Huang (1995), and Lee et al. (1996). These studies offer mixed results – data from other countries frequently, but not exclusively, results in rejections of the random walk model.

Moving to multivariate analysis, it is also interesting to know if a second asset price, \( y_t \), provides any useful information for forecasts of the original asset price, \( x_t \). If additional price information significantly improves a forecast of \( x_{t+1} \), then we conclude that the market is inefficient since this information could be profitably exploited. In Granger’s (1986) words, if two price series come from a jointly efficient, speculative market, they cannot be cointegrated. To explore this issue, I use the Johansen (1988) cointegration procedure and vector autoregression (VAR) analysis to examine if there is any significant relationship between the voucher and equity markets.

Cointegration has been used by many researchers to assess the degree of influence between separate equity markets (Taylor and Tonks, 1989; Kasa, 1992).
1992; Arshanapalli and Doukas, 1993; Palac-McMiken, 1997) and among separate assets traded within one common security market (MacDonald and Power, 1993; Chelley-Steeley and Pentecost, 1994; He, 1997).

The following sections provides sketches of the procedures mentioned above in order to facilitate the interpretation of the results in Section 5. However, since all of the techniques are standard and well-documented, the discussion is kept brief.

3.1. Univariate tests

Since a necessary condition for a random walk is that a price sequence must contain a unit root, I begin the analysis with two different unit root tests. The classic Augmented Dickey–Fuller (ADF) approach uses OLS to estimate the following equation:

\[ D_{xt} = \mu + \phi x_{t-1} + \sum_{i=2}^{k} \gamma_i D_{xt-i} + \epsilon_t, \]

where \( x_t \) is the natural logarithm of the price series, \( \mu \) a constant, \( \epsilon_t \) an error term, and \( \Delta \) indicates the first-difference of the series. Under the null hypothesis, \( \phi = 0 \) implying that the price series is integrated of order 1, I(1).

Due to Perron’s (1989) observation that stationary data with a breaking trend biases ADF tests towards acceptance of the unit root null hypothesis, many recent time-series tests have been designed to accommodate potential breaks. Since the presence of structural change is particularly problematic in data from developing and emerging markets, I also use the Zivot and Andrews (1992) – ZA hereafter – unit root procedure which tests the null hypothesis that the series is I(1) with no break versus the alternative that the data is I(0) with a breaking deterministic trend. Unlike the Perron (1989) procedure, the ZA test does not require a priori knowledge of the break date, avoiding the problems of “pre-testing”. To further investigate the possibility of breaks, I also employ Vogelsang’s (1997) recently developed structural break procedure which tests the null hypothesis of no structural change. It has the useful feature that it can identify the potential date of a regime change regardless of whether the data is stationary or not. Both the ZA and Vogelsang tests start by estimating the following equation:

\[ D_{xt} = \mu + \theta DU_t + \eta t + \delta DT_t + \phi x_{t-1} + \sum_{i=1}^{k} \gamma_i D_{xt-i} + \epsilon_t, \]

13 Also, prices must be I(1), or higher, for there to be any possibility of a cointegration relationship so the results of the unit root tests are relevant to the multivariate analysis that follows.
where $x_t$ is again the log of the price level. Referring to the time of break as $T_b$, the dummy variables in Eq. (2) are defined as: $DU_t = 1$ if $t > T_b$, 0 otherwise, and $DT_t = t - T_b$ if $t > T_b$, 0 otherwise.

The ZA test sequentially estimates (2) for values of $T_b = 2, \ldots, T - 1$, where $T$ is the total number of observations remaining after accounting for first-differencing and lagged terms. The appropriate test statistic is the maximum absolute value of the $t$-statistic for the parameter $\phi$, which is used in an analogous fashion to the Dickey–Fuller $t$-statistic in Eq. (1). The non-standard distributions are available in Zivot and Andrews (1992).

The Vogelsang procedure also proceeds by sequentially estimating (2) for values of $T_b$ that satisfy $0.01T < T_b < 0.99T$. 14 The Sup Wald (also called the SupF) statistic used in this test is the maximum, over all possible $T_b$, of the Wald statistic for testing $\theta = \delta = 0$. The estimated break date is simply the value of $T_b$ associated with the Sup Wald statistic. Appropriate asymptotic distributions for stationary and non-stationary series are tabulated in Vogelsang (1997).

Although the two concepts are often confused, a unit root process is not identical to a random walk process. A random walk must possess a unit root and is therefore a restricted version of the unit root hypothesis. However, while a random walk requires returns to be uncorrelated, a unit root process may have predictable increments depending on the exact distribution of the error process. Furthermore, Campbell et al. (1997) note that unit root procedures are generally not designed to test for unpredictability; rather they focus on the permanence of the shocks. For these reasons, I employ the variance ratio test from Lo and MacKinlay (1988, 1989) which is a direct test of the random walk hypothesis. This procedure exploits the fact that the variance of a random walk must be a linear function of the time interval over which the data is sampled. For instance, if the random walk hypothesis holds, then the variance of data sampled at yearly increments must be 12 times the variance of the monthly differences. Lo and MacKinlay define the following variance ratio (VR) statistic:

$$VR(q) \equiv \frac{\hat{\sigma}_q^2(q)}{\hat{\sigma}_a^2},$$

where $q$ denotes the number of periods in the holding interval, $\hat{\sigma}_a^2$ is the unbiased variance estimator of the one-period returns, and $\hat{\sigma}_q^2(q)$ the unbiased estimator of $1/q$th of the variance of the $q$-period returns. If the sample value of $VR(q)$ is statistically different from 1, then the random walk null is rejected.

14 The truncation of the sample (trimming) is required for the asymptotic results of the test statistic to be nondegenerate. See Vogelsang (1997) for details.
To test this hypothesis, Lo and MacKinlay provide the following two (asymptotically) standard normal test statistics:

\[
\psi(q) \equiv \sqrt{T-1} (VR(q) - 1) \left( \frac{2(2q-1)(q-1)}{3q} \right)^{-1/2} \sim N(0, 1),
\]

\[
\psi^*(q) \equiv \frac{\sqrt{T-1} (VR(q) - 1)}{\hat{\sigma}} \sim N(0, 1),
\]

where in Eq. (5), \( \hat{\sigma} \) is the heteroskedasticity-consistent estimator of the asymptotic variance of \( VR(q) \). \(^{15}\) Hence, \( \psi^*(q) \) is heteroskedastic-robust – an important feature since there is growing evidence in the finance literature of asset-return variance clustering.

### 3.2. Multivariate tests

Since the joint behavior of the Hungarian voucher and equity market is also of interest, I use the Johansen cointegration procedure to examine the possibility of a cointegrating relationship between the two price series. Consider the column vector, \( Z_t \), composed of \( n \), I(1) price series following the (vector) autoregression process,

\[
Z_t = \Omega_1 Z_{t-1} + \Omega_2 Z_{t-2} + \cdots + \Omega_k Z_{t-k} + \mu + \epsilon_t,
\]

where \( \epsilon_t \) is an \( n \times 1 \) vector of Gaussian errors and time is indexed by \( t = 1, \ldots, T \). It is possible to rewrite (6) into the error-correction form,

\[
\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \Gamma_2 \Delta Z_{t-2} + \cdots + \Pi I Z_{t-1} + \mu + \epsilon_t.
\]

Johansen (1988) shows that the rank of the \( n \times n \) matrix, \( \Pi \), equals the number of independent cointegrating relationships. Under the assumed hypothesis that each series in \( Z_t \) is individually non-stationary, the rank of \( \Pi \) must be less than \( n \). In the extreme case where the rank of \( \Pi \) is zero, Eq. (7) simply specifies a standard VAR model in first differences.

There are two test statistics for determining the rank of \( \Pi \), commonly called the trace test (\( \lambda_{\text{trace}} \)) and the maximum eigenvalue test (\( \lambda_{\text{max}} \)), which differ in their maintained alternative hypothesis. The test statistics are:

\[
\lambda_{\text{trace}}(r) = -T \sum_{i=r+1}^{n} \ln(1 - \tilde{\lambda}_i),
\]

\[
\lambda_{\text{max}}(r, r+1) = -T \ln(1 - \tilde{\lambda}_{r+1}),
\]

\(^{15}\) See Lo and MacKinlay (1988, 1989) and Campbell et al. (1997) for details and proofs.
where \( \tilde{\lambda}_{r+1}, \ldots, \tilde{\lambda}_n \) are the \( n - r \) smallest eigenvalues from the estimated \( \Pi \) matrix. The \( \lambda_{\text{trace}} \) statistic is used for testing the null hypothesis that the number of distinct cointegrating vectors is less than or equal to \( r \) against a general alternative. The \( \lambda_{\text{max}} \) statistic tests the null that there are exactly \( r \) versus \( r + 1 \) cointegrating vectors.

4. Description of the data

The analysis is based on daily data from 8 March 1993 through 30 December 1997 yielding a total of 1257 observations spanning nearly 5 years. The starting date was selected based on the availability of consistent daily observations from the voucher market. Voucher prices are daily average Forint (Ft) prices for vouchers purchased on the BSE, and the data comes from Értékpapír és Tozsdé and the Information Center of the BSE. The stock exchange index, BUX, which has a base value of 1000 in January 1991, was obtained from Bloomberg and Datastream. Indices from Bratislava (SAX), Frankfurt (DAX), Prague (PX), Vienna (ATX), and Warsaw (WIG) also come from Datastream.

To summarize the Hungarian stock and voucher data, Figs. 1–4 present plots of stock and voucher prices as well as their daily rates of return. Vouchers were issued with a face value of Ft 1000 which was intended to be its nominal value when it was used to purchase state assets in privatization auctions and sales. In levels, the voucher trading price on the BSE remained stable during 1993 at about 60% of face value. However, after increasing to Ft 785 at the end of November 1993, the prices steadily declined, falling through the Ft 150 level at the end of 1995 only to recover to nearly face value by the end of 1997.

The summary statistics shown in Table 3 describes the series. Particularly noteworthy are the high values of the Ljung–Box \( Q \) statistic for the data which indicates the presence of serial correlation that would be expected from
Fig. 2. Daily compensation voucher prices (March 1993–December 1997).

Fig. 3. Daily stock returns.

Fig. 4. Daily voucher returns.
non-stationary data. Moreover, even the asset returns exhibit high $Q$ values which are indicative of potential autocorrelation.

5. Results

5.1. Univariate tests

Table 4 presents the results of ADF tests for the log and daily return data. The number of lags in the ADF test was determined by using the recommended practice of “testing-down” from a large number of lags. Starting from a maximum of 20 lags, the lags included in Eq. (1) were pared down until the coefficient on the highest remaining lag was significantly different from zero.  

For log data, the low absolute value of the $t$-statistics does not reject the presence of unit roots. The evidence for the daily return data, on the other hand, implies stationarity at the 1% level of significance.

As mentioned previously, one should be cautious when dealing with data that may contain structural breaks since a structural change will bias results towards the non-rejection of unit root. To address these concerns, Table 4 also presents the ZA maximum $t$-statistic described previously. These results support the hypothesis that the data in logs is I(1) with no break versus the alternative of an I(0) series with a breaking deterministic trend.

While the ZA test is designed primarily to detect unit roots in the presence of shifts in the trend function, to test directly for structural change I use the

\[16\text{ As a precaution, the error structure from the ADF regression was checked for autocorrelation using the Ljung–Box } Q\text{ statistic. No significant problems were detected for any of the lag structures selected.} \]
Vogelsang (1997) sup Wald test which can identify breaks regardless of whether the data is stationary or not. The results are displayed in Table 5. The null hypothesis of no structural change is not rejected in either levels or returns so I conclude that the data does not appear to suffer from any major structural breaks. Hence, the preliminary analysis suggests that the data is reasonably well behaved: the data in levels appears to be characterized by a non-breaking unit root process and the daily returns are stationary.

Since a unit root process does not necessarily imply that returns are uncorrelated, we next turn to evidence of uncorrelated increments from the variance ratio test. Table 6 presents results for the Hungarian stock market price index and the voucher price series for various holding periods ranging from 2 to 100 days. 17 For comparison, the table also displays variance ratio coefficients for other regional stock markets and the German DAX index.

Table 4
Unit root tests

<table>
<thead>
<tr>
<th>Series</th>
<th># lags</th>
<th>ADF t-statistic</th>
<th>ZA max t-statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td>Log stock price</td>
<td>19</td>
<td>-0.12</td>
<td>3.48</td>
</tr>
<tr>
<td>Stock returns</td>
<td>18</td>
<td>-6.66*</td>
<td>7.82*</td>
</tr>
<tr>
<td>Log voucher price</td>
<td>11</td>
<td>-1.11</td>
<td>4.23</td>
</tr>
<tr>
<td>Voucher returns</td>
<td>10</td>
<td>-10.00*</td>
<td>10.69*</td>
</tr>
</tbody>
</table>

* Asymptotic critical values for the ADF test at the 1% and 5% significance levels are -3.43 and -2.86, respectively. Source: Hamilton (1994).

Table 5
Vogelsang Sup Wald test for structural change

<table>
<thead>
<tr>
<th>Series</th>
<th># lags</th>
<th>Sup Wald statistic</th>
</tr>
</thead>
<tbody>
<tr>
<td>Log stock price</td>
<td>19</td>
<td>6.38</td>
</tr>
<tr>
<td>Stock returns</td>
<td>18</td>
<td>1.95</td>
</tr>
<tr>
<td>Log voucher price</td>
<td>11</td>
<td>10.44</td>
</tr>
<tr>
<td>Voucher returns</td>
<td>10</td>
<td>1.43</td>
</tr>
</tbody>
</table>

* Asymptotic critical values for the 1%, 5%, and 10% significance levels are 19.90, 15.44, and 13.62 in the stationary case and 30.44, 25.27, and 22.60 for unit root data, respectively. Source: Vogelsang (1997).

17 Because the variance-ratio test based on long-horizon holding periods (relative to the available series) suffers from power loss and bias, the maximum holding period is limited to 100 days. See Campbell et al. (1997) for additional discussion.
The results show a striking pattern. Focusing on the heteroskedasticity-robust test statistic, \( \psi^*(q) \), the Hungarian BUX does not reject the random walk hypothesis for any value of \( q \) examined. Moreover, for holding periods longer than 5 days, the voucher price series also appears to follow a random walk. 18 For both series, the differences between \( \psi(q) \) and \( \psi^*(q) \) appear to be driven primarily by variance clustering since correcting for heteroskedasticity reduces

Table 6
Variance ratios for daily returns

<table>
<thead>
<tr>
<th>Number of days, ( q ), in holding period</th>
<th>2</th>
<th>3</th>
<th>5</th>
<th>10</th>
<th>20</th>
<th>50</th>
<th>100</th>
</tr>
</thead>
<tbody>
<tr>
<td>Hungary VR(( q ))</td>
<td>1.04</td>
<td>1.09</td>
<td>1.12</td>
<td>1.16</td>
<td>1.47</td>
<td>1.75</td>
<td>1.82</td>
</tr>
<tr>
<td>BUX ( \psi(q) )</td>
<td>1.64</td>
<td>2.37*</td>
<td>2.08*</td>
<td>1.71</td>
<td>3.42**</td>
<td>3.34**</td>
<td>2.53*</td>
</tr>
<tr>
<td>( \psi^*(q) )</td>
<td>0.55</td>
<td>0.84</td>
<td>0.81</td>
<td>0.76</td>
<td>1.63</td>
<td>1.84</td>
<td>1.68</td>
</tr>
<tr>
<td>Hungary VR(( q ))</td>
<td>1.23</td>
<td>1.22</td>
<td>1.10</td>
<td>0.98</td>
<td>1.06</td>
<td>1.08</td>
<td>1.21</td>
</tr>
<tr>
<td>Voucher ( \psi(q) )</td>
<td>8.25**</td>
<td>5.18**</td>
<td>1.67</td>
<td>-0.18</td>
<td>0.19</td>
<td>0.37</td>
<td>0.66</td>
</tr>
<tr>
<td>( \psi^*(q) )</td>
<td>3.71**</td>
<td>2.28*</td>
<td>0.74</td>
<td>-0.09</td>
<td>0.10</td>
<td>0.25</td>
<td>0.50</td>
</tr>
<tr>
<td>Austria VR(( q ))</td>
<td>1.07</td>
<td>1.08</td>
<td>1.11</td>
<td>1.12</td>
<td>1.25</td>
<td>1.33</td>
<td>1.37</td>
</tr>
<tr>
<td>ATX ( \psi(q) )</td>
<td>2.61**</td>
<td>2.06*</td>
<td>1.93</td>
<td>1.30</td>
<td>1.81</td>
<td>1.48</td>
<td>1.15</td>
</tr>
<tr>
<td>( \psi^*(q) )</td>
<td>1.31</td>
<td>1.09</td>
<td>1.08</td>
<td>0.79</td>
<td>1.24</td>
<td>1.16</td>
<td>0.97</td>
</tr>
<tr>
<td>Czech Republic PX VR(( q ))</td>
<td>1.35</td>
<td>1.62</td>
<td>1.94</td>
<td>2.14</td>
<td>2.40</td>
<td>3.07</td>
<td>3.36</td>
</tr>
<tr>
<td>( \psi(q) )</td>
<td>10.84**</td>
<td>12.98**</td>
<td>13.42**</td>
<td>10.57**</td>
<td>8.72**</td>
<td>8.02**</td>
<td>6.43**</td>
</tr>
<tr>
<td>( \psi^*(q) )</td>
<td>7.50**</td>
<td>8.57**</td>
<td>8.79**</td>
<td>7.11**</td>
<td>6.09**</td>
<td>6.11**</td>
<td>5.42**</td>
</tr>
<tr>
<td>Germany VR(( q ))</td>
<td>0.96</td>
<td>0.94</td>
<td>0.94</td>
<td>0.84</td>
<td>0.89</td>
<td>0.74</td>
<td>0.77</td>
</tr>
<tr>
<td>DAX ( \psi(q) )</td>
<td>-1.27</td>
<td>-1.47</td>
<td>-1.03</td>
<td>-1.67</td>
<td>-0.76</td>
<td>-1.13</td>
<td>-0.72</td>
</tr>
<tr>
<td>( \psi^*(q) )</td>
<td>-0.64</td>
<td>-0.77</td>
<td>-0.58</td>
<td>-1.02</td>
<td>-0.52</td>
<td>-0.86</td>
<td>-0.59</td>
</tr>
<tr>
<td>Poland VR(( q ))</td>
<td>1.21</td>
<td>1.36</td>
<td>1.54</td>
<td>1.72</td>
<td>2.03</td>
<td>2.46</td>
<td>2.85</td>
</tr>
<tr>
<td>WIG ( \psi(q) )</td>
<td>7.35**</td>
<td>8.74**</td>
<td>8.67**</td>
<td>7.66**</td>
<td>7.40**</td>
<td>6.47**</td>
<td>5.74**</td>
</tr>
<tr>
<td>( \psi^*(q) )</td>
<td>4.96**</td>
<td>5.80**</td>
<td>5.69**</td>
<td>5.07**</td>
<td>5.07**</td>
<td>4.74**</td>
<td>4.40**</td>
</tr>
<tr>
<td>Slovakia VR(( q ))</td>
<td>1.14</td>
<td>1.21</td>
<td>1.47</td>
<td>2.22</td>
<td>3.21</td>
<td>4.53</td>
<td>4.02</td>
</tr>
<tr>
<td>SAX ( \psi(q) )</td>
<td>4.80**</td>
<td>4.94**</td>
<td>7.32**</td>
<td>12.15**</td>
<td>14.91**</td>
<td>14.71**</td>
<td>8.83**</td>
</tr>
<tr>
<td>( \psi^*(q) )</td>
<td>1.78</td>
<td>1.85</td>
<td>2.76**</td>
<td>4.17**</td>
<td>5.29**</td>
<td>6.29**</td>
<td>4.56**</td>
</tr>
</tbody>
</table>

*a Variance-ratio test for the random walk hypothesis. Under the null hypothesis VR(\( q \)) equals 1 and the test statistics \( \psi(q) \) and \( \psi^*(q) \), defined in the text, are asymptotically distributed standard normal. Starting date for sample period is 8 March 1993 except for data from the Czech Republic PX40 which starts on 6 April 1994, and the data from the Slovak Republic SAX which begins on 14 September 1993. Ending date for all data is 30 December 1997.

** Variance ratios are statistically different from 1 at 1% level of significance.

The results show a striking pattern. Focusing on the heteroskedasticity-robust test statistic, \( \psi^*(q) \), the Hungarian BUX does not reject the random walk hypothesis for any value of \( q \) examined. Moreover, for holding periods longer than 5 days, the voucher price series also appears to follow a random walk. 18 For both series, the differences between \( \psi(q) \) and \( \psi^*(q) \) appear to be driven primarily by variance clustering since correcting for heteroskedasticity reduces...
the values of the variance ratio statistic. The low values of $\psi^*(q)$ for the two established markets, Austria and Germany, result in no rejections of a random walk – a conclusion consistent with the notion that these are mature and well-functioning markets.

For nearly all holding periods, the data from the recently established equity markets in the Czech Republic, Poland and Slovakia strongly rejects the random walk hypothesis at a 1% level of significance. The only exception is Slovakia where the heteroskedastic-robust test statistic for 2 and 3 days does not reject the null.

With the exception of Germany, point estimates of the variance ratios are greater than one which indicates positive serial correlation. These findings agree with Lo and MacKinlay (1988) who find variance ratios greater than unity for weekly NYSE–AMEX data and Huang (1995) who reports similar behavior in Asian markets.

One seemingly plausible explanation for positive autocorrelations is the presence of infrequent trading which would cause returns to show positive, but spurious, correlations as discussed in Poterba and Summers (1988). However, while it is possible that the “thinness” in these markets is revealing itself in the variance ratio estimates, there are several reasons to suspect that non-trading effects do not explain all of the bias. Based on a model of non-trading, Lo and MacKinlay (1988) provide evidence that for an equal-weight portfolio to produce the degree of positive autocorrelation implied by the variance ratios in the markets that do not follow a random walk in this sample, over 40% of the securities in each index would have to be untraded in a one week period.\footnote{The 5-day autocorrelation coefficient can be calculated as $VR(10)/VR(5) - 1$ which yields a weekly first order autocorrelation coefficient of approximately 10% for the Czech Republic and Poland and 51% for Slovakia. There is a narrower index for the Warsaw market, the WIG20, which covers a shorter time-span than the WIG. Variance ratios based on this index reject the random walk at the 5% level for all holding periods less than 19 days. The estimated weekly autocorrelation coefficient is 7.8% which still implies that on average, 6 of the 20 securities in the index would have to remain idle in a given week.}

Furthermore, since the PX, WIG, and SAX are value-weighted portfolios, the actual fraction of non-traded shares would have to exceed 40% since the value-weighting tends to mitigate the non-trading bias. Finally, unlike the other more narrowly defined indices, the WIG is actually a comprehensive index, covering all the companies listed on the Warsaw Stock Exchange. If illiquidity was a significant factor, then one would expect the level of autocorrelation in the WIG to be higher than that for the PX and SAX, an assertion which is not supported by the data. For all of these reasons, I conclude that the autocorrelation in the Central European stock markets is not entirely due to non-trading effects.
5.2. Multivariate tests

While the evidence from the univariate analysis largely supports weak-form efficiency in the individual Hungarian equity and voucher markets, there is another dimension to be explored. This is the possibility that one market may contain information that is useful in forming forecasts of the other. If both markets are jointly efficient then one should not be able to detect a cointegrating relationship between the two series, or if a cointegrating vector exists, it should not be economically useful. Table 7 displays estimates of the cointegration test statistics, (9) and (10), estimated by the Johansen maximum likelihood procedure.

It is clear from Table 7 that the null hypothesis of zero cointegrating vectors is not rejected by either the $\lambda_{\text{trace}}$ or the $\lambda_{\text{max}}$ test at any conventional level. Thus we conclude that the rank of $\Pi$ in Eq. (7) is zero and there is no common stochastic trend shared by the voucher and share markets.

Since the lack of a cointegrating relationship implies that Eq. (7) is an (unrestricted) vector autoregression, we next turn to the impulse response plot and tests of Granger causality between the two series. The impulse response plots are displayed in Fig. 5.

Each panel shows the behavior of the return of a particular asset to a one standard deviation innovation in each variable where the dashed lines show two-standard error confidence intervals. This particular figure corresponds to a system where stock returns are placed prior to voucher returns, but interchanging the ordering does not change the results significantly. It is apparent from the figure that an innovation to a particular return has a substantial, but short-lived, effect on the shocked-variable’s own return, disappearing within one or two days. The impulse responses in Fig. 5 also show that the cross-effects of innovations are extremely weak. The rapid speed at which informa-

<table>
<thead>
<tr>
<th>Table 7</th>
</tr>
</thead>
<tbody>
<tr>
<td>Estimates of $\lambda_{\max}$ and $\lambda_{\text{trace}}$ &lt;sup&gt;a&lt;/sup&gt;</td>
</tr>
<tr>
<td>Null hypothesis</td>
</tr>
<tr>
<td>---</td>
</tr>
<tr>
<td>$\lambda_{\text{trace}}$ test</td>
</tr>
<tr>
<td></td>
</tr>
<tr>
<td>$\lambda_{\max}$ test</td>
</tr>
<tr>
<td></td>
</tr>
</tbody>
</table>

<sup>a</sup>Johansen (1988) cointegration statistics. The $\lambda_{\text{trace}}$ statistic tests the null hypothesis of $r$ cointegrating vectors against the alternative of greater than $r$ cointegrating relationships. The $\lambda_{\max}$ statistic tests the null hypothesis of exactly $r$ vs. $r + 1$ cointegrating vectors. Estimation includes a constant in the cointegrating vector and 18 lags as determined by the LR statistic. Critical values are taken from Osterwald-Lenum (1992).
tion is incorporated into asset prices on the BSE is similar to findings for other equity markets. For instance, using VAR techniques, Eun and Shim (1989) report that the response of the NYSE to an innovation is completed within one day and that the Canadian stock market response to shocks in New York are likewise completed in one day.

The informal conclusions from the impulse response exercise are quantified by the results of the Granger causality test shown in Table 8. Clearly there is no detectable Granger causality from stocks to vouchers which implies that knowledge of the past stock price series would not improve forecasts of voucher prices. Interestingly, there does seem to be weak evidence of Granger causality from vouchers to stock prices. As mentioned in

Table 8
Granger causality tests\(^a\)

<table>
<thead>
<tr>
<th>Equation for</th>
<th>F-statistic</th>
<th>Probability value</th>
</tr>
</thead>
<tbody>
<tr>
<td>(\Delta s_t)</td>
<td>1.39</td>
<td>0.13</td>
</tr>
<tr>
<td>(\Delta v_t)</td>
<td>3.96</td>
<td>(5 \times 10^{-8})</td>
</tr>
</tbody>
</table>

\(^a\)Estimated equation is:

\[
\begin{bmatrix}
    \Delta s_t \\
    \Delta v_t
\end{bmatrix} =
\begin{bmatrix}
    A_{10} \\
    A_{20}
\end{bmatrix} +
\begin{bmatrix}
    A_{11}(L) & A_{12}(L) \\
    A_{21}(L) & A_{22}(L)
\end{bmatrix}
\begin{bmatrix}
    \Delta s_{t-1} \\
    \Delta v_{t-1}
\end{bmatrix} +
\begin{bmatrix}
    e_{1t} \\
    e_{2t}
\end{bmatrix},
\]

where \(A_{ij}\) denotes the coefficients in the lag polynomial \(L\) and \(\Delta s_t\) and \(\Delta v_t\) are the returns to stocks and vouchers, respectively. The null hypothesis for the F-test is that all entries of \(A_{12}\) equal zero in the equation for stock returns, and all values of \(A_{21}\) are zero in the equation for voucher returns.

![Fig. 5. Impulse response plots.](image)
Section 2, information regarding the course of privatization was considered to be an important indicator of the government’s economic policies. Hence, it is possible that information regarding privatization plans (as captured by the voucher prices) was also being reflected in the share prices with a lag – an indication of an informational inefficiency.

There is an interesting similarity between the voucher/equity market behavior shown here and the bond/equity market evidence described in Fama and French (1993) and Rahman and Mustafa (1997). It appears that in both cases, factors influential in the bond (or voucher) market affect equity markets but not the other way around. In this sense, this is empirical support for the view that vouchers are a type of government bond.

6. Conclusion

This paper has presented an array of time-series tests exploring the properties of the Hungarian stock market, with a particular emphasis on the interaction between equity and compensation voucher trading.

The univariate evidence shows that the Hungarian equity and voucher markets are remarkably efficient. In particular, variance ratio tests indicate that the random walk model is a fitting description of the equity and voucher markets in Hungary. Furthermore, while stock indices from Western markets follow random walks, other recently-established Central European equity markets consistently reject the random walk hypothesis, implying violations of weak-form efficiency. Thus, in a comparative sense the Hungarian stock market behavior is closer to that of established markets than its peers. Factors that likely contribute to the BSE’s efficiency are the presence of cross-listed firms and the participation of foreign investors. However, since vouchers were traded only in Budapest and foreign ownership of compensation vouchers was explicitly prohibited, these factors do not explain the efficiency of voucher trading. At the same time, it should be pointed out that it is also possible that the variance ratio tests are sensitive to the trading volume and that the indices from Hungary, Germany and Austria are simply composed of firms with higher volume. 20

Based on the multivariate results, there is little indication of significant interaction between voucher and equity markets, despite the fact that vouchers often constituted an important part of the trading volume on the BSE. The Johansen procedure was unable to identify significant cointegrating vectors and the impulse responses from vector autoregressions show little influence

20 Data limitations prevent testing this hypothesis through the construction of indices consisting of firms with similar trading volumes.
either from stock returns to voucher returns or the other way around. Granger causality tests confirm that statistically there is no causation from stocks to vouchers. However, there appears to be some causation in the other direction, behavior which resembles bond and equity market interaction in developed markets.

Overall, the bulk of the evidence points to several conclusions. First, there does not seem to be much detectable linkage between voucher and stock markets in Hungary, except for the extremely weak Granger causality from vouchers to stocks. This observation can be attributed to two possible relationships which cannot be exclusively identified. On one hand, the evidence is consistent with the notion that the pricing of these assets are jointly efficient: knowledge of one does not significantly improve forecasts of the other. On the other hand, another interpretation is that the lack of a significant cointegrating vector implies that there is no long-run equilibrium relationship between voucher and equity prices. This suggests that fundamentals that drive equity returns in Hungary are not affecting the voucher markets. Therefore, it appears that vouchers are deriving their value primarily from exogenous factors such as perceptions about future privatization policies and that these policies have only a limited effect on equity prices. Furthermore, the fact that voucher prices follow a random walk implies that this information was rapidly incorporated into the market. Thus, the Hungarian experience provides evidence that voucher tradability can occur efficiently and may not suffer significantly from major informational asymmetries.

Second, despite the BSE’s immaturity and the limited number of traded securities, the Hungarian stock exchange is extremely efficient at processing information. This has some bearing on the discussion about privatization policies and financial market development. Much of the research at the beginning of the transition expressed skepticism about the role of stock markets because of the belief that early markets would suffer from low liquidity and high volatility (Lipton and Sachs, 1990) or would be unable to cope with the noisy environment (Tirole, 1991). As a result, policy recommendations were biased towards the Western European model of universal banking with significant involvement of commercial banks in the oversight of privatized firms. As measured by market efficiency, Hungary’s experience is inconsistent with this view.

Finally, an additional contribution of this paper is to show that while mass privatization is associated with the rapid formation of tradable securities, the stock markets created as a result of large-scale privatizations – the Czech and Slovak Republics in particular – are not efficient. One obvious interpretation is that the large number of listings created by rapid privatization results in relatively scattered and disbursed trading. This possibility was discussed in The World Bank (1996) World Development Report which notes that trading activity and share prices are lower in rapid privatizers because of low demand
and weak institutions. In this respect, one lesson from Hungary is that the modest number of listed firms and the slower pace of stock market expansion is actually advantageous to efficient trading. Alternatively it may be the case that these younger exchanges require more time to mature and that the Budapest stock market has simply had a longer adjustment period. This question, however, is left for future research.

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References