

Burhop, Carsten; Gelman, Sergey

**Working Paper**

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and expected returns at the  
Berlin Stock Exchange,  
1892-1913

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# **Transaction costs, liquidity and expected returns at the Berlin Stock Exchange, 1892-1913**

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# Transaction costs, liquidity and expected returns at the Berlin Stock Exchange, 1892-1913<sup>\*</sup>

Carsten Burhop<sup>†</sup> / Sergey Gelman<sup>‡</sup>

## Abstract

We estimate effective spreads and round-trip transaction costs at the Berlin Stock Exchange for the period 1892-1913 using daily stock market returns for a sample of 27 stocks. Our results show that transaction costs at the main stock exchange in a bank-based financial system at the turn of the 20<sup>th</sup> century were quite low and about comparable to transaction costs in modern markets. Nonetheless, transaction costs varied substantially over time and across securities, whereby the cross-sectional variation could be substantially explained by firm size and time variation by crises. Furthermore, we find surprising evidence that transaction costs decrease the expected excess returns. Thereby size and momentum premia are of expected signs while market beta has no significant influence on the cross-sectional return variation.

JEL-Classification: G 12, G 14, N 23

Keywords: Transaction Costs; Effective spreads; Economic History; Germany

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† Universität zu Köln, Seminar für Wirtschafts- und Unternehmensgeschichte, Albertus-Magnus-Platz, 50923 Cologne, Germany, and Max-Planck-Institut zur Erforschung von Gemeinschaftsgütern

‡ International College of Economics and Finance, State University – Higher School of Economics, Pokrovskij Bulvar 11, 109028 Moscow, Russia, Email: sgelman@hse.ru

## I. Introduction

High transaction costs on a security market are a sign of market inefficiency. For example, Jensen (1978) demonstrates that transaction costs can limit the ability of traders to profit from mispricing. Moreover, a plethora of papers starting with Amihud and Mendelson (1986) shows that high transaction costs increase the expected return on assets and thus reduce asset prices. Furthermore, Pagano (1989) proves the existence of equilibrium with high transaction costs, low trading activity, and high volatility as well as the existence of a Pareto-superior equilibrium with low transactions costs, high trading activity, and low volatility. In sum, high transaction costs reduce prices and increase volatility. By and large, the empirical literature based on modern data supports this view. For example, Asparouhova et al. (2010), Eleswarapu (1997) and Chalmers and Kadlec (1998) support the hypothesis of a positive risk premium for illiquidity; Baltagi et al. (2006) and Hau (2006) illustrate that higher transaction costs come along with higher volatility.

We measure the size of effective spreads and transaction costs on an early stock market located in a bank-based financial system: the Berlin Stock Exchange between 1892 and 1913. One would expect that effective spreads and transaction costs were comparatively high in this market for at least two reasons: first, technological progress during the last century – e.g., the introduction of electronic trading platforms – should have resulted into lower spreads and costs. Second, stock markets are supposed to be small and inefficient in the traditional German financial system. However, we show that both, effective spreads and transaction costs at the Berlin Stock Exchange a century ago were of about the same size as they are in modern financial markets. Our favourite measure of transaction costs indicates that the cost for a roundtrip transaction were about 0.98 percent of the share price. This compares to an estimate of 1.23 percent for the largest decile of firm listed at the New York Stock Exchange for the period 1963-1990 (Lesmond et al., 1999).

This result fits well with recent empirical literature investigating Germany's early stock market history. According to the traditional view, Germany has had a bank-based financial system dominated by large joint-stock credit banks since the late 19<sup>th</sup> century (see Fohlin, 2007, chapters 1 and 2 for a review). However, recent research points to the significance and efficiency of the German stock market at the turn of the 20<sup>th</sup> century. First of all, Rajan and Zingales (2003) show that the ratio of stock market capitalisation over GDP and the number of listed corporations per million inhabitants was higher in Germany than in the U.S. in 1913. Moreover, the German stock market was not only deep, but also efficient: Weigt (2005: 199) and Baltzer (2006) show that stock price differentials among German stock exchanges and between the Berlin Stock Exchange and other major European stock exchanges were small; stock price differentials were arbitrated away. Furthermore, Weigt (2005: 224) demonstrates that stock prices reflected the risk and return characteristics of the shares quite well. In addition, Gelman and Burhop (2008) illustrate the weak information efficiency of the Berlin Stock Exchange. Furthermore, Gehrig and Fohlin (2006) estimate in a paper closely related to our

work the effective spreads of samples of Berlin traded shares for the benchmark years 1880, 1890, 1900, and 1910.

Qualitatively, our findings are similar to the results presented by Gehrig and Fohlin (2006). We supplement their findings by showing that transaction costs vary substantially over time and across stocks. We can link the variation of transaction costs to crises and to the size of the company in the cross-section. Our findings suggest that transaction costs were particularly high during periods of financial distress, e.g., the 1901 banking crisis and the 1913 war-crisis, and for smaller enterprises (the latter in support of Gehrig and Fohlin (2006)). Moreover, we test whether transaction costs had a positive impact on expected returns, by including a transaction costs measure into CAPM and multifactor models. Surprisingly, we reject the hypothesis: the full transaction costs measure has a negative risk premium, if size is accounted for. Such puzzles have been discovered in previous literature in modern data sets (Chordia et al. 2009, etc.), but have not been fully explained. Investigating this issue in depth seems to be a promising research avenue. Thereby we find – in contrast to Weigt (2005) – that the CAPM does not hold for the period 1892-1913, whereas the size and momentum effects are present.

The remaining parts of the paper are organised as follows. In Section II, we give a short description of the Berlin Stock Exchange at the turn of the 20<sup>th</sup> century and we describe our data sources. Two measures of effective spreads and one measure of round-trip transaction costs are illustrated in Section III, along with brief description of implemented regression techniques. The results are presented in Section IV, followed by our conclusion in Section V.

## **II. Market Structure and Data Description**

During the period of industrialisation, universal banks played a major role in the German economy: they facilitated its ‘take-off’, they provided loans to the modern industrial firms, and they monitored industrial corporations. Close bank-industry relationships ameliorated liquidity constraints of industrial firms and thereby made investment easier and supported industrial growth.<sup>1</sup> Nevertheless, bank loans were only one way to ensure the financing of an industrial enterprise. Self-financing of investments using retained profits was significant in Imperial Germany. Yet, by far the most important source of capital was equity. Share capital accounted for more than half of the assets of Germany’s industrial corporations until World War I (Fohlin, 2007: 174). These equities were issued at one of Germany’s stock exchanges. By the early 1870s, the Berlin Stock Exchange was the main stock exchange in Germany with several hundred listed corporations. Until World War I, the number of Berlin-quoted companies increased to about 1,000.

Shares were traded at the Berlin Stock Exchange six days per week using a call auction mechanism. Official brokers balanced purchase and sale orders and set the final binding price

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1 An excellent review of the literature is provided by Fohlin (2007).

of all orders once a day. This ensured a high level of transparency at the Berlin Stock Exchange throughout the period considered here (Fohlin, 2007: 235). Yet, quoted spreads are not available, making an estimation of effective spreads especially useful. Turning to transactions costs, we may distinguish three types of observable costs: taxes, broker fees, and bank fees. Transactions at German stock exchanges were taxed from 1881 onwards. During this year, a stamp tax on stock market transaction at a fixed rate of 0.20 Mark per transaction was introduced. This tax was transformed into a percentage tax in 1885. More specifically for the period under consideration here, the stock market turnover tax was 0.01 percent of the underlying transaction value between 1892 and April 1894. From May 1894 onwards, the tax was doubled to 0.02 percent; another increase to 0.03 percent followed in October 1900. In addition to turnover taxes, the fees for brokers influence transaction costs. The fee for official brokers (*Kursmakler*) was 0.05 percent of the underlying transaction value, whereas the fee for private brokers (*Privatmakler*) was only 0.025 percent (Gelman and Burhop, 2008). Furthermore, provisions for the banks or other intermediaries varied between 0.1 and 0.33 percent (Weigt, 2005: 192). In sum, broker fees, provisions for intermediaries, and turnover taxes add up to a total cost for a roundtrip transaction (i.e., buying and selling of a share) in the range of 0.252 to 0.82 percent.

To investigate the size of effective spreads and transaction costs and to evaluate whether they changed over time, we use daily stock prices for the period 31 December 1891 to 31 December 1913 collected from the *Berliner Börsenzeitung* – Germany’s leading financial daily of the pre-1913 period – for a sample of 27 continuously traded corporations from the Berlin stock exchange. The data were obtained from Gelman and Burhop (2008) who construct a daily stock market index for the period 1892-1913. The sample contains 6,692 daily returns. The average daily return was 0.023 percent and they display distributional properties similar to modern stock market returns. Descriptive statistics of individual stocks are shown in Table 1.

**Table 1. Distributional properties of stock returns of the Gelman-Burhop-index constituent companies**

	Name	Mean (ann.)	Median (ann.)	Max.	Min.	Std. Dev.	Skew- ness	Kur- tosis	Proportion of zero re- turns	$\rho(1)$	Average Market Cap. (mill M)
1	AG für Anilinfabrikation Allgemeine	0.0678	0.0000	0.1257	-0.2270	0.0082	-3.85	126.25	0.1638	0.0008	2727
2	Elektrizitätsgesellschaft Berlin-Anhaltinische	0.0336	0.0000	0.0526	-0.0611	0.0065	-0.18	11.94	0.0807	0.0820	14997
3	Maschinenbau	0.0134	0.0000	0.1037	-0.0878	0.0078	-0.65	25.62	0.1540	-0.0474	1179
4	Bochumer Bergwerk (Lit C)	0.2457	0.0000	4.6522	-0.3611	0.0603	68.59	5285.9	0.2192	0.0060	351
5	Deutsche Bank	0.0294	0.0000	0.0333	-0.0544	0.0042	-1.64	24.72	0.1001	-0.0119	32778
6	Dresdner Bank	0.0129	0.0000	0.0446	-0.0554	0.0048	-0.85	17.74	0.1062	0.0230	19931
7	Darmstädter Bank (BHI)	0.0006	0.0000	0.0642	-0.0846	0.0044	-1.21	40.81	0.1903	-0.0231	16689
8	Deutsche Jute Spinnerei und Weberei	0.0465	0.0000	0.0674	-0.1040	0.0079	-0.53	16.89	0.2061	0.0024	430
9	Deutsche Spiegelglas	0.0687	0.0000	0.0921	-0.0838	0.0080	-0.30	18.32	0.1877	0.0716	643
10	Erdmannsdorfer Spinnerei	0.0001	0.0000	0.1143	-0.0774	0.0106	0.62	13.70	0.2497	-0.0425	286
11	Gelsenkirchener Bergwerksgesellschaft	0.0254	0.0000	0.0484	-0.0858	0.0071	-0.82	13.87	0.0555	0.0240	15580
12	Gerresheimer Glashütten	0.0480	0.0000	0.0739	-0.1208	0.0079	-1.61	33.11	0.2452	-0.0551	1112
13	Hallesche Maschinenfabriken	0.0266	0.0000	0.1000	-0.2788	0.0093	-6.04	169.13	0.1666	-0.0287	667
14	Harpener Bergbau AG	0.0212	0.0000	0.0668	-0.0682	0.0075	-0.28	11.77	0.0517	0.0282	10400
15	Kattowitzer AG für Bergbau und Eisen	0.0348	0.0000	0.0609	-0.0603	0.0068	-0.26	13.77	0.1313	0.0402	4652
16	Maschinenfabrik Kappel	0.0711	0.0000	0.2014	-0.2138	0.0113	-0.61	48.92	0.1725	0.0721	340
17	Norddeutsche Wollkämmerei	0.0192	0.0000	0.0738	-0.0838	0.0080	0.01	15.22	0.2025	0.0272	1926
18	Schaffhausen'scher Bankverein	0.0000	0.0000	0.0454	-0.0409	0.0037	0.07	22.24	0.2688	0.1121	14043
19	Oberschlesische Portland- Cement AG	0.0390	0.0000	0.1267	-0.0943	0.0098	0.72	23.25	0.1515	-0.0238	440
20	Rheinische Stahlwerke	0.0151	0.0000	0.1095	-0.1427	0.0085	-0.77	39.52	0.1230	-0.0337	3745
21	Rositzer Zuckerfabrik	0.0435	0.0000	0.0833	-0.0826	0.0092	0.05	11.73	0.1413	0.0237	717



22	Chemische Fabrik vormals Schering	0.0162	0.0000	0.0652	-0.0657	0.0083	0.17	10.17	0.1630	0.0610	1298
23	Schlesische Zinkhütten	0.0336	0.0000	0.1079	-0.0853	0.0066	-0.52	33.99	0.2360	-0.0582	7947
24	Schlesische Leinen-Industrie	-0.0043	0.0000	0.0703	-0.0679	0.0063	-0.56	21.28	0.3194	-0.1558	1015
25	Schultheiss Brauerei	0.0097	0.0000	0.0897	-0.0853	0.0057	-1.23	65.22	0.1714	-0.2016	2522
26	Siemens Glas-Industrie	0.0287	0.0000	0.0438	-0.0576	0.0058	-1.12	19.17	0.1966	-0.0273	2290
27	Stettiner Chamottewaren	-0.0096	0.0000	0.0804	-0.2538	0.0088	-3.85	126.25	0.1452	0.0281	2640
	Gelman-Burhop index	0.0687	0.0909	0.0296	-0.0562	0.0032	-1.68	30.78	n/a	0.165	161344
	Equally-weighted price index	0.0327	0.0600	0.1780	-0.0325	0.0036	17.80	913.90	n/a	0.167	161344

Notes: Mean and median returns are presented on the annual basis (mean daily return x300) for illustrative purposes.

### III. Econometric technique

In an information-efficient stock market, prices of stocks should fluctuate randomly. However on the real-world stock exchanges the presence of transaction costs induces some predictive patterns, i. e., deviations from random walk. Uncovering these patterns and analyzing them allows tracking back effective spreads or even full transaction costs.

Several straightforward measures are employed to assess the size of effective spreads, which is the difference between the ask price and the bid price and constitutes a part of overall transaction costs. As the Berlin Stock Exchange was an auction and not a quote market there are no quoted spreads available. However there are grounds to assume that several large players (e. g., banks) did systematically submit buy and sell orders simultaneously, thus acting as de-facto market makers. Therefore indirect measures of effective spreads are here of particular value, since they constitute the only way to assess the differences between ask and bid prices.

The simplest measure, introduced by Roll (1984), rests upon the assumption that liquidity-caused trades (“initiated other than by new information”) are equally probable to happen at ask and bid prices. In absence of strong movements of the “true” value of stock, which is midpoint of bid and ask prices, this assumption leads to negative autocorrelation of observed stock returns. Moreover, the degree of autocorrelation strictly depends on the spread size. More formally, let  $r_{i,t} = (p_{i,t} - p_{i,t-1})/p_{i,t-1}$  be the return on stock  $i$  in period  $t$ , where  $p$  denotes the stock price. Then the effective spread  $S_i^R$  is given by equation (1):

$$(1) \quad S_i^R = 2\sqrt{-\text{cov}(r_{i,t}, r_{i,t-1})}$$

However, a brief glance at the descriptive statistics tells you that the daily returns of about half of the stocks in the Berlin Stock Exchange index were – to a small degree – positively autocorrelated. As a result, the Roll measure yields negative spreads in numerous cases and is generally downward biased (George et al., 1991).<sup>1</sup> Obviously, spreads were not the only source of autocorrelation. Thus, George et al. (1991) develop a measure of effective spreads, which eliminates one of the non-microstructural sources of autocorrelation – positive autocorrelation of time-varying expected returns. This GKN measure of effective spreads starts from the return of an equally-weighted market portfolio of  $N$  securities, i.e.,  $r_t^m = 1/N \sum_{i=1}^N r_{i,t}$ . Then, to account for

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1 If the autocorrelation is positive, then the expression under the root becomes undefined in real numbers. Here we follow the trick applied in the literature (e. g. Gehrig and Fohlin 2006, Roll 1984) – we simply pull the minus out of the root, thus defining the Roll ratio as:

$$S_i^R = \begin{cases} 2\sqrt{-\text{cov}(r_{i,t}, r_{i,t-1})} & \text{if } \text{cov}(r_{i,t}, r_{i,t-1}) \leq 0 \\ -2\sqrt{\text{cov}(r_{i,t}, r_{i,t-1})} & \text{if } \text{cov}(r_{i,t}, r_{i,t-1}) > 0 \end{cases}$$

Since the negative spreads do not have economic sense, the estimated negative Roll effective spread measures can be interpreted as the degree of inappropriateness of this measure to estimate spreads (for further discussion see Campbell et al. 1997, p. 135).

autocorrelation of expected returns, a regression of the return of stock  $i$  on the past return of the market portfolio is run. The covariance of residuals  $\hat{\varepsilon}_{i,t}$  of such a regression – see equations (2) and (3) – can be interpreted as a measure of effective spreads.<sup>2</sup>

$$(2) \quad r_{it} = \hat{\alpha}_i + \hat{\beta}_i r_{t-1}^m + \hat{\varepsilon}_{it}$$

$$(3) \quad S_i^{\text{GKN}} = 2\sqrt{-\text{cov}(\hat{\varepsilon}_{it}, \hat{\varepsilon}_{it-1})}$$

One should note that the Roll measure of effective spreads and the GKN measure of effective spreads are identical whenever the regression parameter  $\beta_i = 0$ . If, however,  $\beta_i > 0$ , then both measures of effective spreads are downwards biased, but the bias of the GKN measure is smaller compared to Roll's measure.<sup>3</sup>

A measure of transaction costs was proposed by Lesmond et al. (1999). The LOT measure reflects the total costs of a roundtrip transaction, which includes, opposite to Roll and the GKN measures, not only the difference between bid and ask prices, but also all further expenses carried by the trader, including the price change induced by the trade itself (so called price impact, see Lesmond 2005). In contrast to the Roll and GKN measures, the LOT measure has an information-economics theoretical foundation, whereas Roll and GKN are based on an liquidity argument. The LOT measure is based on the idea that transactions will only occur if the deviation of the market price from the true value of a stock is larger than the costs of a transaction. Thus, there are upper and lower thresholds –  $\tau_i^l$  and  $\tau_i^h$  – such that the measured return  $r$  is non-zero only if the true return exceeds the threshold:

$$(4a) \quad r_{i,t} = R_{it}^* - \tau_i^l \text{ if } r_{i,t}^* < \tau_i^l$$

$$(4b) \quad r_{i,t} = 0 \text{ if } \tau_i^l < r_{i,t}^* < \tau_i^h$$

$$(4c) \quad r_{i,t} = R_{it}^* - \tau_i^h \text{ if } r_{i,t}^* > \tau_i^h.$$

Thereby the true return depends on the market return  $r_{m,t}$  in a linear way:  $r_{i,t}^* = \beta_i r_{m,t} + e_{i,t}$ .

The estimated difference between the upper and the lower threshold – i.e.  $\tau_i^h$  less  $\tau_i^l$  – is a measure of the roundtrip transaction costs.

We use the following maximum likelihood estimator, developed by Lesmond et al. (1999), to estimate the LOT measure:

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2 Here we again encounter the problem of undefined spreads (in real numbers) if the residual covariance in (3) is positive. We address it in the similar way, as by the Roll measure (see fn. 3).

3 The GKN measure is unbiased if and only if the forecasted market return is a perfect proxy for the expected return of stock  $i$ .

$$\begin{aligned}
L(\tau_i^l, \tau_i^h, \beta_i, \sigma_i | r_{it}, r_{mt}) &= \prod_1 \frac{1}{\sigma_i} \phi \left[ \frac{r_{it} + \tau_i^l - \beta_i r_{mt}}{\sigma_i} \right] \times \prod_0 \left[ \Phi \left( \frac{\tau_i^h - \beta_i r_{mt}}{\sigma_i} \right) - \Phi \left( \frac{\tau_i^l - \beta_i r_{mt}}{\sigma_i} \right) \right] \\
(5) \times \prod_2 \frac{1}{\sigma_i} \phi \left[ \frac{r_{it} + \tau_i^h - \beta_i r_{mt}}{\sigma_i} \right] \\
\text{S.T. } \tau_i^l &\leq 0, \tau_i^h \geq 0, \beta_i \geq 0, \sigma_i \geq 0,
\end{aligned}$$

Where  $\Phi()$  is the standard normal cumulative distribution function. Thereby region 1 (indicated by “1” below the  $\Pi$ ) corresponds to the negative expected latent variable when the observed is nonzero ( $\hat{r}_{it}^* < 0$ , or equivalently  $r_{mt} < 0$  and  $r_{it} \neq 0$ ), region 2 – to the positive expected latent variable if observed is nonzero ( $r_{mt} > 0$  and  $r_{it} \neq 0$ ) and region 0 corresponds to the observation with zero observed returns ( $r_{it} = 0$ ).  $\sigma_i$  denotes the root out of the residual variance, measured over non-zero returns region.

In contrast to the Roll and GKN measure, the LOT measure thus includes the bid-ask spread, commissions, transaction taxes, costs of information acquirement and processing as well as price impact. It is our measure of transaction costs. Its size should be larger than the Roll and the GKN measure and it should be larger than the regulated costs, i.e., the sum of broker fees, provisions, and transaction taxes.

As we find considerable differences in transaction costs across companies, similar to Gehring and Fohlin (2006), we run cross-section regressions of estimated transaction costs on a set of explanatory variables:

$$(6) \bar{S}_i = \alpha + \beta' X_i + \varepsilon_i,$$

where  $X$  denotes a vector of explanatory variables and  $\beta$  a vector of corresponding coefficients. However, as we observe remarkable time variation of transaction cost estimates we also run a panel regression:

$$(7) S_{it} = \alpha + \beta' X_{it} + \mu_i + \lambda_t + v_{it},$$

where  $\mu_i$  denote cross-sectional individual effects,  $\lambda_t$  denote period effects and  $v_{it}$  is an idiosyncratic error term.

We rely on the standard technique in the asset pricing literature, the Fama-MacBeth (1973) regression, when analyzing the impact of transaction costs on the cross-sectional variation of returns. It is based on the assumption, that expected returns of stocks are fully described by the linear combination of risk premia and factor loadings for all relevant factors:

$$E[Z_i] = \lambda' B_i,$$

whereby  $Z_{it} = r_{it} - r_{ft}$  denotes excess return,  $\lambda'$  is a transposed vector of risk-premia and  $B_i$  is a vector of factor loadings or risk characteristics of company  $i$ . Given the values of factor loadings

for each stock in each period the risk premia are estimated running  $T$  cross-section regressions (one for each period) and averaging the estimates:

$$Z_{it} = \lambda_t' B_{it}$$

$$\bar{\lambda} = \frac{1}{T} \sum_{t=1}^T \hat{\lambda}_t$$

The corresponding standard errors for each  $k$ -th element of the risk-premia vector are calculated from the corrected time variance of the estimated premia:

$$\text{var}[\lambda_{kt}] = \frac{1}{T} \sum_{t=1}^T (\hat{\lambda}_{kt} - \bar{\lambda}_k)^2$$

$$\text{stderr}[\bar{\lambda}_k] = \sqrt{\text{var}[\lambda_{kt}] \cdot \left(1 + \frac{\bar{z}_{mt}^2}{\text{var}[z_{mt}]}\right) / T},$$

where  $z_{mt}$  denotes the excess return of the market index. For the risk factor  $k$  to be priced the corresponding risk premium should be significantly different from zero.

## IV. Results

Table 2 presents the annual averages over all trading days and all shares for each measure as well as the average for the full sample period 1892-1913; Figure 2 shows the evolution of transaction costs on a daily frequency.

As expected, the Roll-measure of effective spreads underestimates the transaction costs at the Berlin Stock Exchange. The Roll-measure points towards negative effective spreads for more than half of the years in the sample and for the whole sample period on average. This finding is similar to the size of the Roll measure reported by Gehrig and Fohlin (2006) for the four benchmark years 1880, 1890, 1900, and 1910. According to them, the Roll measure varies between -0.06 percent (in 1900) and 0.10 percent (in 1910). The average Roll measure calculated by Gehrig and Fohlin (2006) is -0.025 percent, quite close to the average value of -0.029 percent calculated from our data. From an economic point of view, negative spreads make no sense. Consequently, we investigate if the GKN measure provides more sensible results.

The GKN measure performs slightly better. On average, effective spreads were only 0.062 percent of the share price. Thus, our estimate of the GKN effective spread is even lower than the average value calculated by Gehrig and Fohlin (2006) of 0.179 percent. In Figure 2 it can be seen that GKN spread is lower than the Roll measure for a short period of 1895, suggesting (temporary) negative autocorrelation of time-varying expected returns. Moreover, one should note that the GKN effective spread is negative in seven years. All these results convey that de-facto market making did not influence the stock price dynamics.

**TABLE 2: ANNUAL AVERAGE OF TRANSACTION COSTS**

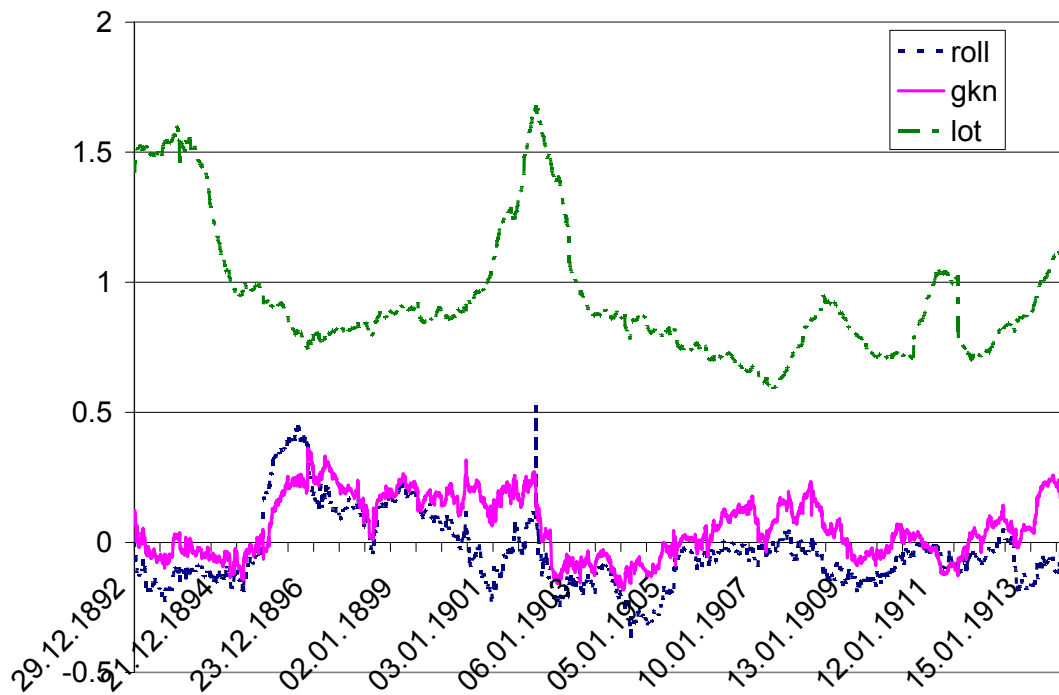
Year	Roll	GKN	LOT
1892	-0.015	0.088	1.454
1893	-0.104	0.003	1.584
1894	-0.152	-0.058	1.072
1895	0.214	-0.015	0.925
1896	0.190	0.254	0.805
1897	0.158	0.195	0.814
1898	0.210	0.231	0.908
1899	0.097	0.198	0.878
1900	-0.221	0.050	1.029
1901	0.124	0.215	1.678
1902	-0.133	-0.074	0.977
1903	-0.233	-0.155	0.848
1904	-0.198	-0.043	0.825
1905	-0.054	0.083	0.696
1906	-0.052	0.030	0.658
1907	0.015	0.148	0.775
1908	-0.156	0.011	0.846
1909	-0.115	-0.024	0.731
1910	-0.008	-0.038	1.039
1911	-0.072	0.028	0.713
1912	-0.143	0.049	0.883
1913	0.015	0.185	1.124
Average	-0.029	0.062	0.966

*Own calculations based on daily returns for 27 stocks for the period 1892-1913.*

*Expressed in percent of share price, equally weighted averages. Two outliers were dropped.*

Our favourite measure, the LOT measure of transaction costs, is positive for every year between 1892 and 1913. According to this measure, transaction costs at the Berlin Stock Exchange varied between 0.66 percent (in 1906) and 1.68 percent (in 1901). The transaction costs were positive for any randomly chosen yearly period and always higher than the lower bound of the regulated fees. The average transaction costs amounted to 0.97 percent. Therefore, we broadly confirm the result presented by Gehrig and Fohlin (2006), who estimated an average LOT measure of 0.71 percent for the four benchmark years 1880, 1890, 1900, and 1910.

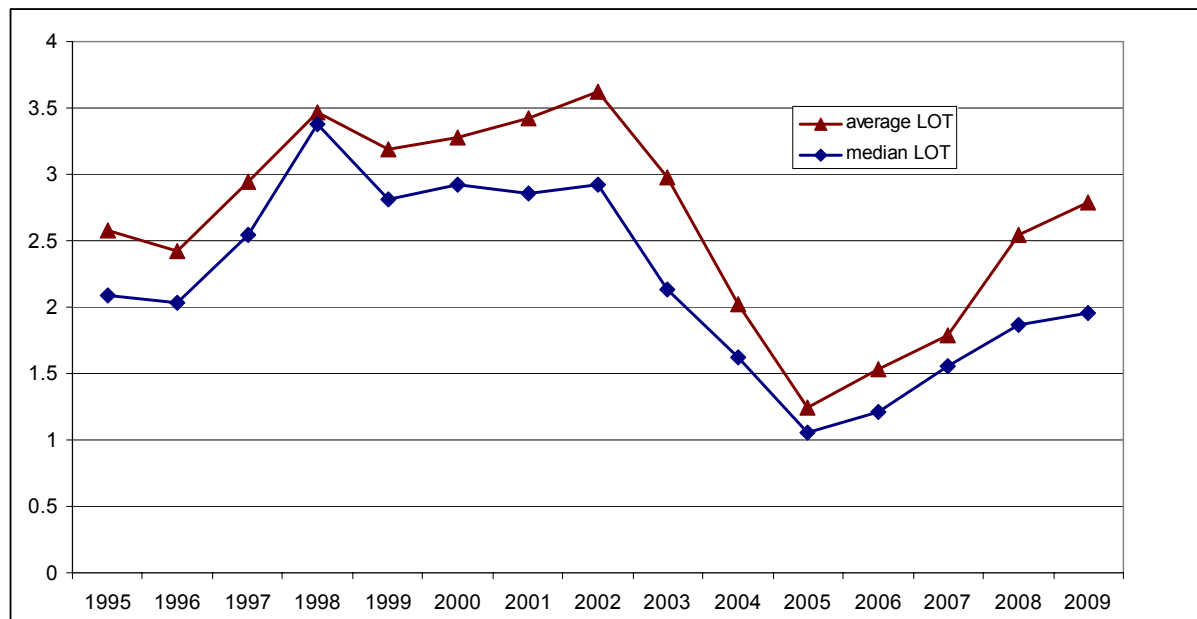
**Figure 2. Rolling estimates of transaction costs**



*Notes:* Estimates of transaction cost measures according to Roll (1984), equation (1), GKN (1991) – equations (2)-(3), and LOT (1999) – equations (4a)-(4c) for the subsample of the fixed length of 1 year (305 trading days). Dates on X-axes mark the end of the yearly period. Equally weighted averages over all 27 stocks constituting the Gelman and Burhop (2008) index.

Compared to modern data, it may come as a surprise that transaction costs were rather stable at the German stock exchange over the last century. We find that the 27 companies under study at the turn of the twentieth century had, on average, lower transaction costs than the 2<sup>nd</sup> tier German blue chips at the turn of the twenty-first century: Applying the same technique to 47 MDAX companies for 1995-2009 yields average and median LOT measures of 2.6 percent and 2.1 percent. Interestingly, we observe that the annual average transaction costs for the crisis years 1998 and 2000/2001 are with 3.3-3.5 percent clearly above the overall sample mean.

**Figure 3. Annual LOT measures for MDAX companies**



*Notes:* Estimates of transaction cost measures according to LOT (1999) – equations (4a)-(4c) for each year. Equally weighted averages and medians over 47 stocks belonging to MDAX index.

Evidence for other modern stock markets supports the impression that transactions costs were quite low at the Berlin Stock Exchange already a century ago. Goyenko et al. (2009) document LOT measures for Dow Jones Industrial Average index of 0.6 percent in mid 1970s and 1980s which is comparable to our results for the Berlin Stock Exchange index in mid 1900s. Very advantageous is the comparison to the modern emerging markets: stocks in the Gelman-Burhop (2008) index have lower transaction costs according to the LOT measure than any of the 31 emerging markets in the 1990s, covered in the study of Lesmond (2005): their average transaction costs range from 2.3 percent for Taiwan to 18 percent for Russia. Some, but not all, of this reported superiority of Berlin Stock Exchange at the turn of the 20<sup>th</sup> century can be explained by broader sample of securities pro country, analyzed in Lesmond (2005). Furthermore, it seems that transaction costs declined in emerging markets since the late 1990s: Lagoarde-Segot (2009) estimates LOT ratios of less than 0.5 percent for most developing countries for the period 1996-2007.

Additional insights provide inter-company differences in transaction costs (see Appendix 1). Whereas textile companies, such as Deutsche Jute Spinnerei und Weberei and Erdmannsdorfer Spinnerei report LOT measures of (1.1 percent) and (1.7 percent) which could be found also for median Chinese stock (compare Lesmond 2005), the transaction costs of banking sector stocks like Deutsche Bank (0.38 percent) and Dresdner Bank (0.45 percent) is on the same level with Dow Jones companies in 1980s and 1990s (compare Goyenko et al. 2009). These deviations however, cannot be attributed fully to industrial differentiation: companies included into the index stemming from the banking sector have much higher market capitalization, e.g. the value of Deutsche Bank was on average 114 times the value of Erdmannsdorfer Spinnerei.



The explanation may rather have informational origins: assuming the same share of trading relative to market capitalization across companies, the volume of trade for large companies was higher, allowing market makers to incorporate new information. Furthermore, large companies had probably better analyst coverage, providing more thorough information to investors, thus decreasing information asymmetry and providing for lower transaction costs.

Some evidence for this hypothesis can be obtained from a simple cross section regression of average transaction costs on the log of the market capitalization. However, one should be cautious as Amihud and Mendelson (1986) reveal the possibility of a reversed causal relationship: transaction costs can raise expected returns and thus reduce the market capitalization of a company. To avoid the endogeneity problem and to ensure the pre-determinacy we use the market capitalization of 1892 (which is measured at the beginning of the year) to explain company transaction costs averaged over the twenty-two year sample. For the twenty-six companies we obtain (standard errors of estimates are in parenthesis):<sup>4</sup>

$$\overline{S_i^{LOT}} = 4.92 - \underset{(0.62)}{0.19} \cdot \ln \left( MC_i^{1892} \right) + \hat{e}_i, \quad (8)$$

$$R^2 = 0.64, \quad \hat{e}_i \approx (0, 0.20^2)$$

All estimates are highly significant and support the hypothesis that market capitalization decreases transaction costs: raising the market capitalization by 2.3 bln Mark (what corresponds to a one unit change of log market capitalization at the mean of the variable) leads to a 0.19 percentage points narrower spread. Interestingly, our coefficient coincides with the one reported by Gehrig and Fohlin (2006) for the year 1900 for the log of the book value and is considerably close to their results for 1890 and 1910. Thereby market capitalization explains almost two thirds of the inter-company transaction cost variation in our sample.

The question is, whether it is firm-size per se or informational coverage that matters for the transaction costs of stocks. A variable which is not related to size but to informational coverage is what we call the effective dividend forecast error (EDFE). This error can be tracked back from the price change on the first day a stock is traded ex-Dividend and the true value of the dividend payment. The peculiarity of the Berlin stock exchange at that time was that the shareholder meeting, on which the amount of dividend was decided on, was held after the ex-Dividend day. Thus, investors had to forecast the dividend payment when calculating the price adjustment for the loss of the dividend claim. If investors would have perfect foresight (and there would be no news concerning the future long-term performance), in the absence of market frictions the price should

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4 We exclude Bochumer Bergwerk henceforth from the analysis, as it has unusually high transaction costs due to several months long periods of non-trading.

fall exactly by the dividend amount; any misestimation of dividends should yield deviations of price change from the dividend amount.<sup>5</sup>

We assume that for stocks with intensive information coverage such forecasts should be more accurate than for stocks less frequently analyzed in media and therefore this effective forecast error can serve as proxy of information coverage. We hypothesize that lower information coverage (larger absolute ED FE) causes higher spreads. To exclude the possibility of reverse causality we use effective dividend forecast errors for 1892. Extending correspondingly the above regression we obtain (standard errors of estimates are in parenthesis):

$$\overline{S_i^{LOT}} = \underset{(0.66)}{5.08} - \underset{(0.03)}{0.20} \cdot \ln(MC_i^{1892}) - \underset{(2.99)}{0.06} |EDFE_i^{1892}| + \hat{e}_i, \quad (9)$$

$$R^2 = 0.66, \quad \hat{e}_i \approx (0, 0.20^2)$$

The coefficient in front of the information coverage proxy is insignificant and of the wrong sign. Possibly, given that *EDFE* could be very noisy, using the first year results could be insufficient to uncover the relationship.

Therefore, we run regressions of transaction costs on market capitalization and on both market capitalization and *EDFE* in a balanced panel set-up with individual effects. We assume that trade volume is proportional to market capitalization not only across companies but also across time. If higher trade volume of larger firms is associated with lower transactions costs, then we should find the same relationship in the panel regression as in the cross section regressions (8) and (9).

However, as market capitalization is clearly non-stationary over the 22 year sample, we use the fraction of the overall market capitalization contributed by each company. Therefore, we have to include time-effects to account for changes in the overall market capitalization.

To address the endogeneity problem we use alternative specifications of contemporaneous and lagged absolute *EDFE*. Since market capitalization is reported for the beginning of each year, we do not face possible reverse causality.

We perform random effects and fixed effects estimations. Following the result of the Hausman test fixed effects are chosen.

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5 Here we also account for the usance of the Berlin Stock Exchange called Stueckzinsen (see Gelman and Bur-

hop 2008). The full formula is therefore 
$$EDFE = \frac{P_{t^d} - P_{t^d-1}}{P_{t^d-1}} - \frac{0.04}{P_{t^d-1}} + \frac{d}{P_{t^d-1}},$$
 where  $t^d$  is the date of dividend coupon detachment,  $0.04/P_{t^d-1}$  is the Stueckzinsen correction and  $d$  is the actually paid dividend

**Table 3. Panel regression results**

	(1)	(2)	(3)	(4)
Constant	1.01*** (0.07)	0.98*** (0.07)	0.95*** (0.07)	0.97*** (0.05)
$MC_i/\Sigma MC$	-3.22* (1.76)	-3.25 (2.02)	-3.46** (1.61)	-2.76** (1.28)
$t_{1901}$		0.29*** (0.08)		
$t_{1913}$		0.26*** (0.08)		
$ EDFE _{it}$			2.96*** (1.08)	
$ EDFE _{it-1}$				-0.13 (1.02)
Time effects	Y	N	Y	Y
$R^2$	0.56	0.41	0.58	0.60

Estimates of LS fixed effects models for the transaction costs (LOT measures) for the sample period from 1892 to 1913 for the panel of 26 companies of the type:  $S_{it} = \alpha + \beta'X_{it} + \mu_i (+\lambda_t) + v_{it}$ . Standard errors are reported in parenthesis. Values marked with \*\*\*, \*\* and \* are significant at 1%, 5% and 10% level respectively.  $R^2$  is calculated as one minus the fraction of the residual variance to the variance of the dependent variable.

Column 1 in Table 3 supports the size hypothesis: negative coefficients in the regressions with fixed effects indicate that if the share of market capitalization in the index increases by one standard deviation (5 percentage points) transaction costs decrease by 0.16 percentage points. Some loss in statistical significance – now the impact of market capitalization is significant only on the 10% level opposed to the 1% level in cross-section – is probably due to a much larger amount of noise in the company-year level data.

Furthermore, time effects play a substantial role: they explain about 17 percent of the transaction costs variation. About one-eighth of the explanatory power of time can be attributed to two years: 1901 and 1913 (see column 2) – which are known to have caused worsening information efficiency (Gelman and Burhop 2008). In 1901 the bankruptcy of Leipziger Bank, one of Germany's largest banks, caused a stock exchange turmoil and possibly high degree of uncertainty about fair prices of shares which forced speculative traders to act more conservative, thus increasing transaction costs. In 1913, the fear of a Balkan war led to similar effects on the financial market. About 16 percent is explained by cross-sectional individual effects. The loss of the explanatory power of market capitalization (from 64% in the cross-section to about 23% in the

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amount. Here we abstain from accounting for market movements, as many companies have ex-dividend dates on the same day leaving the rest of the market portfolio under-represented.

panel) can be explained by (1) company-specific features, correlated with size and captured by individual effects in the panel set-up and (2) the higher level of noise.

Including the absolute effective dividend forecast error seems to significantly improve the regression fit (see column 3), as we obtain a highly significant and positive relationship, supporting the hypothesis of higher spreads under lower accuracy (and hence information coverage). However, results with lagged absolute EDFE in column 4 disavow this statement, as no significant relationship is found. Since information coverage should not change dramatically from year to year, we argue that the uncovered strong relationship may have the reverse causal direction due to the proxy design. In fact, following the logic of LOT (equations 4a and 4c), investors knowing the exact amount of the future dividend payment would adjust the price by this amount less the transaction cost. Thus, the proxy captures not only forecast accuracy but also market frictions, and with increase of transaction costs absolute EDFE also rises. So we can not find support for our information coverage hypothesis. But the good news conveyed by this interrelation is that the LOT measure seems to be adequate in spite of not accounting for company-specific information.

Our main result – negative correlation of transaction costs with size – proves to be rather stable over time: An equivalent to column 1 panel regression of transaction costs of 47 MDAX stocks over 1999-2009 on the fraction of overall market capitalization yields a coefficient of -5.34, which is also significant on the 10% level and explains about 6% of the variation of transaction costs.<sup>6</sup>

The interdependence of size, information coverage and transaction costs should provide for the presence of the so-called lead-lag effect: the investors of large-cap companies should react on new market-wide information quicker than the investors of small-cap enterprises. Therefore, the returns of large-cap stocks should lead the returns of the small-caps. To test for the presence of this effect we calculate equally-weighted returns of two portfolios: one containing the nine most capitalized companies ( $r_{Lt}$ ) and another one consisting of the nine companies with the smallest market capitalization ( $r_{St}$ ) as of the beginning of 1892. Then we run a simple VAR for the whole period and test the Granger causality, similar to Hou (2007):

$$r_{Lt} = a_{01} + a_{11}r_{Lt-1} + a_{12}r_{St-1} + e_{1t}$$

$$r_{St} = a_{02} + a_{21}r_{Lt-1} + a_{22}r_{St-1} + e_{2t}$$

In order to find empirical support of the effect we should be able to reject  $H_0: a_{21}=0$  and sustain  $H_0: a_{12}=0$ . The estimation yielded following results:

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6 Results available on demand.

$$r_{Lt} = \underset{(4.1 \cdot 10^{-5})}{6.1 \cdot 10^{-5}} + \underset{(0.012)}{0.138} \cdot r_{Lt-1} + \underset{(0.006)}{0.006} \cdot r_{St-1} + e_{1t}$$

$$r_{St} = \underset{(5.2 \cdot 10^{-5})}{9.5 \cdot 10^{-5}} + \underset{(0.016)}{0.271} \cdot r_{Lt-1} + \underset{(0.007)}{0.026} \cdot r_{St-1} + e_{2t} \cdot$$

The null hypothesis that returns on large stocks do not Granger cause returns on smallest companies can be rejected on any conventional significance level, whereas the probability that returns on small stocks do not improve the forecast of large stock returns is  $p=0.28$ . Thereby the previous day returns on large stocks explain about five percent of the variance of the small stock portfolio, whereas the previous returns of small stock do not influence the contemporary returns of large stocks.

A trivial explanation would be a larger portion of non-trading among smaller stocks. In fact, there are more zero daily returns of the constituents of small cap portfolio (however, the data available does not allow to distinguish whether the stock was not traded or traded at the same price to the previous day). But this difference is by far not sufficient to explain the lead-lag relationship: if this cross-autocorrelation would be induced by non-synchronous trading only the ratio of cross-auto-covariances should be equal to the ratio of non-trading probabilities. In our example that would mean about 2.6 times higher non-trading probability for small cap stocks than for the large caps, whereby it is on average only 1.33 times higher. Thus we can hypothesize that the returns on larger stocks lead the returns on smaller stocks, to a substantial extent due to market frictions other than non-trading (such as higher transaction costs). This also means that the stock prices of small companies react with delay to the new information.

The large dispersion of transaction costs should be revealed in asset pricing. As Amihud and Mendelson (1986) noted, given the set of investment opportunities, investors should avoid assets which have lower liquidity yielding same returns. This should in the long run decrease the price of such securities and raise their return. Therefore, in the long run one should find a positive relation between transaction costs and expected returns in the cross-section.

Here we analyze excess returns, calculated as total returns (price changes plus dividends) less the risk free rate. Including dividends is important as the companies may compensate investors with higher dividends for lower prices. In line with asset pricing literature, we use monthly data. We run Fama-MacBeth (1973) regressions for the traditional CAPM and several multifactor extensions, including transaction costs, and other popular factors – namely size and momentum.

As the risk free rate proxy we use the money market rate obtained from NBER (series: 13018). Size is the log of market capitalization and varies on the annual basis. Transaction costs are our LOT estimates, effective spreads are our Roll (1984) estimates, both of them also vary yearly. Market betas and momentum, the latter calculated as the first-order autocorrelation coefficient of daily price percentage changes, are constant for each company throughout the sample. We also include a constant as we do not demean the explanatory variables.

**Table 4. Results of cross-section regressions for the whole sample 1/1892-12/1913**

	(1)	(2)	(3)	(4)	(5)
Constant	.0018 (.0014)	.0030 (.0021)	.0376*** (.0102)	.0399*** (.0104)	.0325*** (.0106)
Market beta $\bar{\lambda}_\beta$	-.0003 (.0019)	-.0010 (.0021)	.00002 (.0021)	-.0016 (.0023)	.0003 (.0020)
Transaction cost $\bar{\lambda}_{TC}$		-.0497 (.1406)	-.4600** (.1833)	-.4563** (.1842)	-.3499* (.2065)
Size $\bar{\lambda}_S$			-.0015*** (.0004)	-.0015*** (.0004)	-.0013*** (.0004)
Momentum $\bar{\lambda}_M$				.0125** (.0059)	
Effective spread (Roll) $\bar{\lambda}_{ES}$					-.1039 (.1338)
Average $R^2$	0.07	0.13	0.17	0.21	0.24
# of stocks	26	26	26	26	26
# of cross-sections	264	264	264	264	264

Estimates of the Fama-MacBeth (1973) regressions for the sample period from 1892 to 1913 for 26 companies. Reported coefficient values  $\bar{\lambda}_k$  are averages of 264 regression estimates of the type:  $Z_{it} = \alpha_i + \lambda_i' B_{it} + u_i$ , where  $\lambda_i'$  denotes the transposed vector of risk premia and  $B_{it}$  denotes the vector of risk factor loadings, which serve as explanatory variables in each cross section. Standard errors are calculated as

$$\sqrt{\text{var}[\lambda_{kt}] \cdot \left(1 + \frac{\bar{z}_{mt}^2}{\text{var}[z_{mt}]}\right) / 264},$$

according to the Fama-MacBeth (1973) procedure with Shanken (1992) correction, and are reported in parenthesis. Values marked with \*\*\*, \*\* and \* are significant at 1%, 5% and 10% level respectively. Average  $R^2$  is an arithmetic mean of  $R^2$  for each cross-section.

Our results in Table 4 reveal that CAPM does not hold for our sample in 1892-1913: the market risk premium is insignificant in all five specifications; moreover it is often of a wrong sign.

The premium for the transaction costs in a two-factor model (column 2) is insignificant. However, adding size results in obtaining significant coefficients for transaction costs, which is in all these three specifications (columns 3-5) negative. This means that for the given size companies with higher transaction costs yield lower average returns. This finding clearly contradicts Ami-

hud and Mendelson (1986) and the major bulk of the empirical literature on this issue. However, similar results have been reported by Chordia et al. (2009: 3657) for the Roll-Gibbs measure of NYSE/AMEX transaction costs in 1976-2002, Brennan and Subrahmanyam (1996) encounter a negative risk premium for proportional spreads and Eleswarapu and Reinganum (1993) report negative relationship of transaction costs and expected returns of NYSE stocks in non-January months of 1961-1990. Brennan and Subrahmanyam (1996) hypothesize that transaction costs proxy some risk factor related to the price level or size, and this relation could cause such an effect. The magnitude of the effect is also economically significant: 0.5% growth of transaction costs (what corresponds to approximately one standard deviation of a LOT measure in cross section) results in 0.23% lower monthly returns (or about 2.76% lower annualized returns).

All three specifications containing size (columns 3-5) provide strong support for the small size premium, as all three coefficients are significant on the 1% level and negative. Thus, according to column 3, companies with market capitalization of 0.8 bln Mark smaller than average yield annualized expected returns 1.8% (0.15% on monthly basis) higher than average. In line with our earlier argumentation, investors holding stocks of small cap companies underlie risk of lower liquidity and scarcer information coverage and require therefore higher returns.

The risk premium for momentum is significant and positive (see column 4): stocks, which daily price percentage changes are positively autocorrelated have on average higher expected returns. As this autocorrelation is negatively related with the Roll (1984) effective spread measure (see Eq. 1), it could be also interpreted as negative risk premium for effective spread size. To test this interpretation directly we include in column 5 Roll measure instead of autocorrelation coefficient. While we obtain the expected sign, the estimate becomes insignificant. Obviously, the transformations – taking square root and scaling by the standard deviation – weaken the interdependence. That leaves space for other explanations of the momentum premium.

Summarizing our study of impact of transaction costs on asset pricing, we find significant negative influence of round-trip transaction costs if the size is controlled for, which is a counter-theoretical and puzzling result. However, similar results have been encountered in earlier research (e.g. Chordia et al. 2009: 3657). Further analysis of this issue, involving use of other transaction costs measures, which do not include price impact, as well as use of further stock characteristics, is necessary. Thereby market risk premium is insignificant, size and momentum premia have expected signs and are significant at least on the 5% level.

## V. Conclusion

Transaction costs at the Berlin Stock Exchange averaged about one percent between 1892 and 1913 according to the measure proposed by Lesmond et al. (1999). Thus, transaction costs a century ago were quite similar to today's cost. This may be surprising in the light of the traditional view of Germany as a bank-based financial system with underdeveloped security markets. However, recent empirical research shows that the German stock market was relatively deep during

the pre-1913 period (Rajan and Zingales, 2003), weakly information efficient (Gelman and Burhop, 2008), and integrated with other German and European financial markets (Weigt, 2005; Baltzer, 2006). Consequently, low transaction costs – which were already reported by Weigt (2005: 198) and Gehrig and Fohlin (2006) – fit well into the picture of an efficient security market during Germany's industrialisation.

We discover also a high variation of transaction costs. The latter can be explained by size of the stock and period uncertainty: we find a negative relation between market capitalization and estimated transactions costs (supporting Gehrig and Fohlin 2006) as well as higher transaction costs in the crisis years 1901 and 1913.

Studying the impact of transaction costs on the cross-section of expected returns we find a significant negative relationship. This result is inconsistent with economic theory and rather puzzling, but similar phenomena have been reported in previous empirical work for different samples (e.g. Eleswarapu and Reinganum 1993 for NYSE stocks 1961-1990). Further research steps, such as including further transaction cost measures and control variables, are needed to shed more light on this puzzle.



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## Appendix 1: Average transaction costs of corporations, included in the investigation

Number	Name	Average Roll measure	Average GKN measure	Average LOT measure
1	AG für Anilinfabrikation	0.050	0.117	0.943
2	Allgemeine Elektrizitätsgesellschaft	-0.233	-0.183	0.520
3	Berlin-Anhaltinische Maschinenbau	0.091	0.172	0.902
4	Bochumer Bergwerk (Lit C)	-0.308	-0.162	3.164
5	Bank für Handel und Industrie	0.001	0.059	0.543
6	Deutsche Bank	0.004	0.029	0.384
7	Dresdner Bank	-0.101	-0.014	0.446
8	Deutsche Jute Spinnerei und Weberei	0.053	0.124	1.109
9	Deutsche Spiegelglas	-0.185	-0.101	1.097
10	Erdmannsdorfer Spinnerei	0.320	0.364	1.689
11	Gelsenkirchener Bergwerksgesellschaft	-0.156	-0.101	0.427
12	Gerresheimer Glashütten	0.168	0.309	1.284
13	Hallesche Maschinenfabriken	0.136	0.258	1.112
14	Harpener Bergbau AG	-0.116	-0.060	0.425
15	Kattowitzer AG für Bergbau und Eisen	-0.121	0.020	0.667
16	Maschinenfabrik Kappel	-0.416	-0.328	1.239
17	Norddeutsche Wollkämmerei	-0.130	-0.007	1.135
18	Oberschlesische Portland-Cement AG	0.038	0.228	1.094
19	Rheinische Stahlwerke	-0.069	0.179	0.781
20	Rositzer Zuckerfabrik	-0.123	-0.052	1.053
21	Schaaffhausen'scher Bankverein	-0.156	-0.071	0.572
22	Chemische Fabrik vormals Schering	-0.291	-0.154	1.001
23	Schlesische Zinkhütten	0.204	0.251	0.959
24	Schlesische Leinen-Industrie	0.427	0.446	1.183
25	Schultheiss Brauerei	0.202	0.217	0.684
26	Siemens Glas-Industrie	0.081	0.162	0.776
27	Stettiner Chamottewaren	-0.147	-0.033	0.905

Source: Gelman and Burhop (2008), own calculations

## Appendix 2. Results of cross-section regressions for the sub-sample 1/1892-12/1901

	(1)	(2)	(3)	(4)
Constant	.0066*** (.0020)	.0072** (.0029)	.0315** (.0140)	.0401*** (.0143)
Market beta $\bar{\lambda}_\beta$	-.0015 (.0029)	-.0015 (.0032)	-.0003 (.0033)	-.0024 (.0035)
Transaction cost $\bar{\lambda}_{TC}$		-.0522 (.1769)	-.2397 (.1999)	-.3396 (.2111)
Size $\bar{\lambda}_S$			-.0011* (.0006)	-.0014** (.0006)
Momentum $\bar{\lambda}_M$				.0161*** (.0062)
Average $R^2$	0.06	0.12	0.17	0.20
# of stocks	26	26	26	26
# of cross-sections	120	120	120	120

Estimates of the Fama-MacBeth (1973) regressions for the sample period from 1892 to 1901 for 26 companies. Reported coefficient values  $\bar{\lambda}_k$  are averages of 120 regression estimates of the type:  $Z_{it} = \alpha_t + \lambda'_t B_{it} + u_i$ , where  $\lambda'_t$  denotes the transposed vector of risk premia and  $B_{it}$  denotes the vector of risk factor loadings, which serve as explanatory variables in each cross section. Standard errors are calculated as

$$\sqrt{\text{var}[\lambda_{kt}] \cdot \left(1 + \frac{\bar{z}_{mt}^2}{\text{var}[z_{mt}]}\right)} / 120,$$

according to the Fama-MacBeth (1973) procedure with Shanken (1992) correction, and are reported in parenthesis. Values marked with \*\*\*, \*\* and \* are significant at 1%, 5% and 10% level respectively. Average  $R^2$  is an arithmetic mean of  $R^2$  for each cross-section.