DETERMINANTS OF COUNTRY BETA RISK IN POLAND

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Abstract

In the paper we analyze determinants of the capital market beta risk in Poland in the monthly period 1996-2002. The beta risk is measured as a time-varying parameter estimated in a regression of the Warsaw stock indexes (WIG and WIG20 separately) on major foreign stock market indexes (DJIA, NASDAQ, DAX and FTSE). The individual monthly beta parameters time series are computed as structural regression parameters estimated for daily data in monthly sub-periods in regressions for WIG and WIG20 indexes on individual foreign stock market indexes. The beta risk is an average of monthly individual beta parameters. We put forward a hypothesis that the estimated beta risk depends on monetary and real variables expressing the economic performance of the Polish economy. Hence, we build monetary and real factors models. As explanatory variables of risk, we examine: income, productivity, trade balance, budget deficit, interest rate and the zloty exchange rate. The risk factors are expressed as differentials relative to the world economy for which stands the U.S. economy. According to Fair and Shiller (1990), we test for relative one-period-ahead predictive performance of monetary and real factors models of capital market risk in Poland in the period 1999-2002. We find that monetary variables as exchange rate and interest rate have relatively more power than real variables in explaining the beta market risk in Poland.

JEL Classification: C2, C5, E6, F3, G1.

Keywords: country beta risk, capital market, risk modelling, econometric model, forecasting, Poland.

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Introduction

Globalization of world's markets and markets of Central and Eastern countries (CEEC) had a considerable influence on their integration. Most of financial decision-making in international setting needs to apply a framework of estimating a country-level risk. This approach is particularly important in an assessment of investment projects, both portfolio and foreign direct, on emerging markets (see e.g. Godfrey and Espinosa, 1996). CEEC financial markets re-emerged in 1990s with the advent of transition from a planned to a market economy. The new markets became a very important factor of the economy restructure and played a prominent role in the process of privatization.

In Poland the Warsaw Stock Exchange (WSE) was opened in April, 1991. Initially, only five companies were listed with trade once a week during the first year of operating the WSE. The situation changed dramatically over twelve years of transition and ongoing process of globalization and integration of world's capital markets. At the end of 2002, stocks of more than 200 companies were listed with a capitalization of over 110 bln PLN (ca. 28 bln U.S. dollars). In the development of capital market in Poland a crucial role was played by a privatization process, inflow of foreign direct investment, development of banking industry, investment funds, and an active role of insurance companies and pension funds. The process of capital market integration in Poland will be further strengthened by the accession of Poland into the EU.

In this paper we aim at studying macroeconomic factors influencing the capital market risk in Poland. We develop an economic model of country beta model risk and search for monetary and real factors that influence asset returns in Poland. Risk is one of fundamental factors that are considered while making assessment of investment projects. There is a large body of literature on examining risk at the country level both for developed and developing countries. Many economists explore the area of beta risk determinants from the political, economic, and financial point of view. Below we give an overview of recent empirical developments in the literature.

We start with a brief overview of political risk influence on asset returns. A comprehensive study has been proposed by Diamonte et al. (1996) who have shown an influential role of political risk on stock returns in emerging and developed markets. They documented a convergence in political risk across countries and found that changes to political risk were more influential on emerging market returns than on developed market returns. While this role of political risk in emerging markets is more pronounced, Diamonte et al. concluded that if global political risk continue to converge, the effects differential between emerging and developed markets may narrow. As a result it shows that macroeconomic factors do become more influential as far as country risk is concerned. Another look at influence of political risk on asset returns has been given by De Haan et al. (1997). They estimated a probit model of country risk, measured as a chance of debt rescheduling, and found little support for political risk to influence the country risk measure but not the influence of economic variables. Consistently with a literature, De Haan et al. suggested that changes to political situation are already discounted in macroeconomic aggregates. An influence of a broad range of different risk measures, both political and economic, on expected asset returns has been also investigated by Erb et al. (1996). They found that risk indexes are highly correlated with fundamental financial attributes and that financial risk variables are more pronounced in explaining future expected asset returns than political risk measures. According to Erb et al., impact of economic and financial risk is most strongly evidenced in the developed markets, while political risk measure helps to some extent in explaining asset returns in emerging equity markets.

Another stream of studies is focused on economic factors of capital market risk. Chang and Pinegar (1987) documented, in accordance with Fama (1981) and Geske and Roll (1983), a negative relationship between stock returns and inflation which varies systematically with securities risk. This effect becomes more negative, the higher increase of securities risk. Another example is Erb et al. (1994). They modelled correlations between equity markets of G-7 countries as functions of financial variables and found that the correlations are influenced by the business cycle. They also found that the correlations were higher when countries were in a common recession, than during recoveries and when countries were out of business cycles phase. The correlations, according to Erb et al., are not symmetric, i.e. they are much higher when markets downgrade. Choi and Rajan (1997) based their analysis on APT model, initiated originally by Ross (1976) and further augmented with macroeconomic variables by Chen et al. (1986). The model included an exchange rate risk as a factor under the assumptions that exchange rate changes are not purely monetary phenomenon and that they influence asset returns due to various real factors influencing deviations from purchasing power parity. Choi and Rajan have found both a positive and a negative impact of exchange rate risk on asset returns in seven major countries excluding the U.S. Groenewold and Fraser (1997), similarly to Choi and Rajan, have tested the macro-factor APT model. They evidenced an influence of short-term interest rate, the inflation rate and the money growth rate on securities returns in Australia. They documented that the APT model is superior to the most widely used CAPM model (originated by Sharpe, 1964) in within-sample tests but the models perform poorly out of sample. In their model variables such as exchange rates, balance of payments, output or employment had less significant impact on asset returns. Brooks et al. (1997) examined the stability of market model betas of U.S. banking industry stocks. They focused on beta stability within the framework of different stages of the banking regulatory process. Brooks et al. have found that regulatory changes influence the stability of beta risk of banks. They also found a similar pattern for non-banks suggesting that the impact on the banking industry is driving the rest of the economy. Bracker and Koch (1999) discussed empirically evolution of global capital market integration within the framework of changing structure of correlation matrix of returns across national equity markets. They modelled potential macroeconomic determinants of the estimated correlation structure and employed the empirical model to generate out-of-sample forecasts compared to non-theoretical models. They indicated significant changes in the correlation matrix of returns both in the short and long run which gives insight to mixed evidence on the stability of the correlation structure. They also applied Dickey-Fuller tests on correlation time series and found that almost all time series contain no unit root. Bracker and Koch have found their economic model to be superior to non-theoretical models as measured by forecast performance. They evidenced that e.g. exchange rate volatility, term structure differentials and real interest differentials across countries have a dampening effect on correlation structure. Gangemi et al. (2000) developed an economic model of the country beta risk in the Australian context. They modelled country betas as a function of macroeconomic variables. The set of variables in their study have been determined in a similar manner as those in e.g. Abell and Krueger, 1989; Bekaert et al., 1996; Erb et al., 1996; Groenewold and Fraser, 1997. The outcome of the paper by Gangemi et al. is that only the trade-weighted exchange rate index had a significant influence on country betas and asset returns. Their results suggested that an appreciation of the home currency has a positive impact on the country beta in Australia and that external shocks play an important role in macroeconomic performance. We also point out the work by Goldberg and Veitch (2002) who developed an economic model of country beta risk in the case of Argentina in the spirit of work by Gangemi et al. (2000) and Erb et al. (1996). They studied the importance of contagion effects of trading partners exchange rate risk on the beta risk of the country operating under a fixed exchange rate regime. They have found that the only economic variables that matter for variations in country beta of Argentina are exchange rates of its trading partners, i.e. Brazil and Mexico.

As we have presented in the literature overview, in empirical research many risk factors of stock returns can be specified, e.g. political, financial, and economic, as well as different risk measures of financial assets can be applied, e.g. variance, semi-variance of returns or conditional variance in GARCH models. The beta risk is an alternative measure of risk.

The aim of our paper, motivated by the literature, was to assess the risk of capital market in Poland within a framework of the market model of beta risk. Poland is an emerging small open economy with strong influences from European and world financial markets. The ongoing transition process from a planned to a market economy offers more and more stable economic environment and investment opportunities. The market performance will be further strengthened by the accession of Poland into the European Union.

We employ the beta risk in an international setting to capture the riskiness of the capital market in Poland. We obtained explicitly time-varying country beta risk measures. We observe time-varying country betas not only in emerging, but also in developed markets, since economic factors capture the existence of business cycles[3].

There are several contributions of our paper. First, we explicitly estimated time-varying beta parameters and used the time series of beta risk as a dependent variable in our model. Second, our motivation was to use relations of home to foreign variables to capture for differentials affecting the Polish economy. Third, we used a procedure of checking out-of-sample predictive quality of our economic models to search for monetary and real factors affecting the country risk. And finally, we have applied the methodology to an emerging market as Poland.

We tend to provide a macroeconomic analysis of country risk factors of monetary and real side origin. The beta risks are regressed on monetary and real variables to test for monetary and real factors that partially influence the capital market risk. The set of macroeconomic variables is generally similar to that used in the literature (see Abell and Krueger, 1989; Bekaert et al. 1996; Erb

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[3] The effects of business cycles on financial risk was studied by e.g. Fama and French (1989); Ferson and Harvey (1991); McQueen and Roley (1993); Erb et al. (1994); Jagannathan and Wang (1996).

et al., 1996; Groenewold and Fraser, 1997; Gangemi et al., 2000). The set of variables included interest rates, nominal exchange rate, income, productivity, trade balance deficit, and a budget deficit. The variables potentially influencing the risk are expressed as home variables related to foreign variables that is somewhat exploratory in nature, given the existing literature. The choice of set of variables is arbitrary and our motivation was to select variables that closely represent the economic performance of the Polish economy. In this paper we aim at extending the existing literature on country beta risk by applying a procedure of forecasting quality test proposed by Fair and Shiller (1990) to search for monetary and real determinants of capital market risk in Poland.

The remainder of the paper is structured as follows. In Section 1, we present a methodology of measurement of country beta risk. Section 2 includes an empirical analysis of monetary and real factors models of beta risk. And finally, in Section 3 we give concluding remarks.

1. Country beta risk: a methodology of measurement

In this section we describe a methodology of measurement of the capital market risk in Poland. We estimate monthly models of risk. We have estimated the beta risk using the market model of beta risk given as (expressed in changes of logs $\cong$ returns):

$$\Delta \log(y_{it}) = \alpha_{ij} + \beta_{ij} \Delta \log(X_{ijt}) + \varepsilon_{it} , \qquad (1)$$

where:

y_i - i th index of the Warsaw Stock Exchange (WSE) (points),

X_{ij} - j th index of foreign stock market for i th index of WSE (points),

ε_i - error term, $\varepsilon_t \sim IN(0, \sigma^2_{\varepsilon_i})$,

$i = \{WIG, WIG20\}$[4],

$j = \{DJIA, NASDAQ, DAX, FTSE\}$.

[4] WIG is the main index of Warsaw Stock Exchange, and $WIG20$ is an index of 20 biggest companies.

To estimate beta risk we have used daily returns close-to-close on WIG and $WIG20$ indexes as well as on foreign indexes in the period January 1, 1996 – December 31, 2002. The sample has been divided into 84 monthly sub-periods. For each monthly sub-period we have estimated parameters α_{ij} and β_{ij} of equation (1). In turn, we obtained eight time series of α and β parameters, i.e. four in the case of returns on index WIG and four in the case of returns on index $WIG20$. Then we focused only on parameters β and calculated an average for each monthly sub-period for indexes WIG and $WIG20$, respectively. Finally, we obtained two time series of average monthly point estimates of parameters β_{WIG} and β_{WIG20}. The beta parameters have been subsequently used as risk measures of capital market in Poland. Below in Table 1 we present descriptive statistics and Jarque-Bera normality test statistics and ADF unit root test statistics for time series of β_{WIG} and β_{WIG20} in full sample and sub-samples.

Table 1. Statistics of variables β_{WIG} and β_{WIG20} - full sample (84 observations)

	β_{WIG}	β_{WIG20}
Average	0.27	0.36
Standard deviation	0.32	0.35
Median	0.26	0.38
Maximum	1.06	1.07
Minimum	-0.62	-0.61
Asymmetry	0.13	-0.28
Kurtosis	3.25	3.21
Jarque-Bera test statistic	0.44 [prob 0.80]	1.22 [prob 0.54]
(A)DF test statistic	-4.39	-4.52

Source: own calculations.

Table 2. Statistics of variables β_{WIG} and β_{WIG20} - sub-sample 1996, M1 – 1999, M1 (37 observations)

	β_{WIG}	β_{WIG20}
Average	0.35	0.41
Standard deviation	0.42	0.45
Median	0.39	0.40
Maximum	1.06	1.07
Minimum	-0.62	-0.61
Asymmetry	-0.35	-0.40
Kurtosis	2.37	2.46
Jarque-Bera test statistic	1.37 [prob 0.50]	1.43 [prob 0.49]
(A)DF test statistic	-5.66	-5.40

Source: own calculations.

Table 3. Statistics of variables β_{WIG} and β_{WIG20} - sub-sample 1999, M2 – 2002, M12 (47 observations)

	β_{WIG}	β_{WIG20}
Average	0.20	0.33
Standard deviation	0.18	0.25
Median	0.19	0.37
Maximum	0.62	0.76
Minimum	-0.25	-0.36
Asymmetry	-0.08	-0.64
Kurtosis	2.59	3.13
Jarque-Bera test statistic	0.38 [prob 0.83]	3.27 [prob 0.19]
(A)DF test statistic	-3.26	-5.34

Source: own calculations.

Based on descriptive statistics of β_{WIG} and β_{WIG20}, we conclude that the capital market in Poland was characterized by a relatively small beta risk with respect to world capital markets. On average β_{WIG} was 0.27 in a full sample while β_{WIG20}, 0.36. This result is consistent with a literature that emerging markets have lower betas than developed markets[5]. On average the betas β_{WIG} and β_{WIG20} are positive which gives a positive correlation of Polish capital market returns with world capital markets.

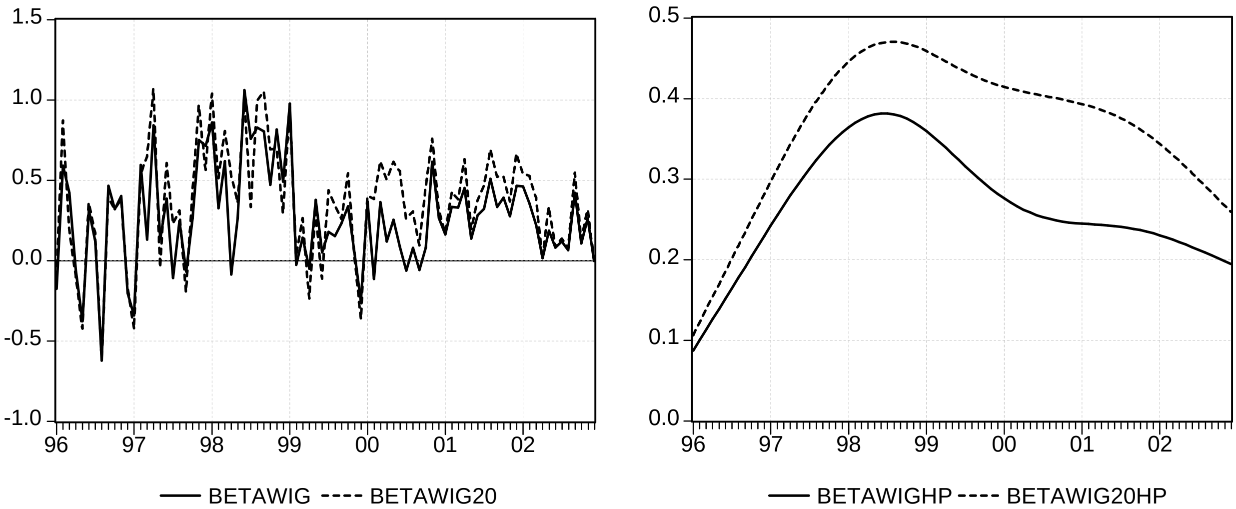
We can also notice that median is close to average in both cases. We can see that β_{WIG} has a positive coefficient of asymmetry which gives a 'fat tail' on the right hand side of the distribution. In the case of β_{WIG20} the coefficient of asymmetry is negative which gives a 'fat tail' on the left-hand side of the distribution. We also have calculated Jarque-Bera normality test statistic (see Jarque and Bera, 1980). For β_{WIG} it is equal to (0.44[0.80])[6] which means we do not reject a hypothesis of normality, while for β_{WIG20} the test statistic is (1.22[0.54]) where we notice larger departures from normality but again we do not reject the null of normality. We also have found that time-varying betas are stationary in full sample and sub-samples by applying the (A)DF test for unit roots. We also have calculated descriptive statistics in sub-periods in which we have noticed different behavior of beta series.

Now let us have a look at plots of variables β_{WIG} and β_{WIG20} (see Fig. 1). In order to see a long-run trend in the data we have smoothed the series by Hodrick-Prescott (HP) filter. We have noticed that both indexes during 1996-98 were characterized by an upward trend. For this period, the time trend slope coefficient for β_{WIG} is equal to 0.0215 and for β_{WIG20} to 0.0226. Both coefficients are statistically significant. The time coefficients for both beta series in the years 1999 – 2002 and in a full sample are statistically insignificant. We conclude that in the first sub-period, i.e. during 1996-98, the capital market in Poland was characterized by an increasing risk. We suggest that this upward trend was also associated with the contagion effects of Asian crisis of 1997 and Russian crisis of 1998. In the second sub-period, according to HP plot, the tendency reversed and betas started to decline. The downward tendency in betas, i.e. declining risk, can be explained by further

[5] See e.g. Harvey (1995) and Erb et al. (1996) who have shown that emerging markets have lower betas with respect to the world market portfolio than developed markets.

development of the capital market and e.g. establishment of Open-end Pension Funds. The Funds are restricted by law to invest in low risk portfolios. The Funds are investing with a high capital relatively to the WSE capitalization and they prevent their portfolios to downgrade and decrease in value. In turn, the demand of the Funds prevents the WSE against sharp declines. Another factor explaining a downward trend in risk during 1999-2002 could be declines in foreign markets, as evidenced by NASDAQ and DAX. This resulted in capital inflow into emerging markets. In consequence, we could observe a lower reaction of Polish indexes that lowered betas and thus the risk in relation to foreign markets.[7] A downward shift in betas can be also associated with a shift in exchange rate regime in Poland from a target zone into free floating after inflation targeting policy was announced by the central bank in 1999 and after introduction of the euro.

Figure 1. Plots of original and filtered variables β_{WIG} and β_{WIG20}



Source: own calculations.

In Section 2 we use the series of β_{WIG} and β_{WIG20} as measures of country beta risk. We propose models of risk with explanatory variables explaining the monetary and real effects of the Polish economy.

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[6] Respective probabilities are given in brackets after test statistics.
[7] There exists an extensive literature which documents that individual stock and portfolio betas are time varying. This is evidenced in e.g. Fabozzi and Francis (1978); Sunder (1980); Alexander and Benson (1982); Bos and Newbold (1984); Faff et al. (1992); Brooks et al. (1992). In the case of Poland see e.g. Wdowinski and Wrzesinski (2003).

2. Modelling market beta risk with monetary and real economy variables

In this Section we present estimation results of modelling beta risk with the use of variables explaining the behavior of the Polish economy. We assume that variables β_{WIG} and β_{WIG20} depend on monetary and real economy variables. As monetary factors we use interest rates and exchange rates. As real factors we use income, labor productivity, trade balance, and budget deficit. As average betas reflect the dependence of the Polish market on foreign markets, we applied a modelling framework in which we have taken relations of Polish variables to foreign variables. We have proxied foreign variables by variables reflecting the U.S. economy. We assumed the following working hypotheses regarding the influence of explanatory variables on beta risk. In the case of monetary variables, we assumed that an increase of interest rate should increase risk as well. We assumed that an increase of interest rate reflects anticipation of inflation growth. In the case of an emerging market like the Polish one, growth of inflation is generally negatively perceived by financial markets as a danger for stable and sustained growth. In emerging economies or economies shifting from a central planning to a market economy prices are influenced by supply shocks and their changes are not of monetary origin to a large extent. That is why inflation growth is transferred to financial markets as a negative signal, i.e. it raises beta risk. With respect to exchange rates we assume that in the short run devaluation gives rise to exports growth and thus to lowering of trade balance deficit. Those effects dominate over price growth due to devaluation. In turn, the exchange rate growth should lower beta risk. This influence was evidenced by e.g. Choi and Rajan (1997); Bracker and Koch (1999); Gangemi et al. (2000); Goldberg and Veitch (2002). In the case of real factors of risk, we assumed that growth of the trade balance deficit and budget deficit reflects the deterioration of the competitiveness of the Polish economy and, in turn, it will increase beta. On the contrary, income and productivity growth compared to the world economy leads to rise in competitiveness and should lower financial risk and stabilize the capital market.

We have splitted our monthly sample into two sub-periods, i.e. 1996, M1 – 1999, M1 (Sample I) and 1999, M2 – 2002, M12 (Sample II) and estimated the models in the sub-periods and in a full sample. The sub-periods have been selected on the basis of data analysis given in Section 1 where we analyzed the tendency of β_{WIG} and β_{WIG20} series. By splitting the sample we wanted to check if estimates are robust to the sample choice. In Section 3 we also forecast beta risk and test out-of-sample forecasting quality of alternative risk models to determine the factors, both nominal and real, most affecting the risk of capital market in Poland. Then splitting the sample serves our forecasting exercise as well.

We have determined many factors potentially influencing the beta risk. Initially broad specification of monetary and real models including foreign exchange rates, inflation, unemployment, and wages has been empirically tested and thus narrowed. We have selected the models that are preferred by their economic and statistical performance. In Tables 4 and 5 we present a summary of estimation results. The estimated models and data used are given in Appendix. In Table 4 we present results of monetary models for β_{WIG} and β_{WIG20}.

Table 4. Models of β_{WIG} and β_{WIG20} for monetary variables

Monetary model (M)																
WIG																
intercept	interest rate	exchange rate	S_e	JB	DW	BG	ARCH	White	Chow	(A)DF	R^2 (adj.)	TP	sample	obs.	equation	forecasting model
0,50 *6,10*	0,44 *1,33*	-0,10 *-2,38*	0,39	0,88 *0,64*	1,67	0,31 *0,58*	2,14 *0,14*	3,15 *0,68*	1,30 *0,30*	-6,25	0,16	53,8%	1996, M05 1999, M01	33	1	
-0,05 *-0,33*	0,02 *1,79*	-0,01 *-1,44*	0,18	0,48 *0,79*	1,98	0,05 *0,82*	0,18 *0,67*	2,62 *0,76*	0,50 *0,95*	-7,18	0,08	45,7%	1999, M02 2002, M12	47	2	
0,05 *0,55*	0,02 *2,19*	-0,03 *-2,24*	0,30	1,28 *0,53*	1,64	1,62 *0,20*	0,05 *0,82*	8,47 *0,13*	0,30 *1,00*	-4,54	0,08	50,8%	1996, M05 2002, M12	80	3	M
WIG20																
-0,32 *-0,53*	0,05 *1,41*	-0,09 *-1,91*	0,43	0,60 *0,74*	1,63	0,38 *0,54*	1,71 *0,19*	3,76 *0,58*	0,84 *0,62*	-6,15	0,10	54,2%	1996, M05 1999, M01	33	4	
-0,16 *-0,95*	0,05 *3,11*	-0,02 *-1,99*	0,22	2,21 *0,33*	1,96	0,04 *0,85*	3,28 *0,07*	4,74 *0,45*	0,37 *0,99*	-7,15	0,19	42,9%	1999, M02 2002, M12	47	5	
0,42 *10,90*	0,21 *2,08*	-0,03 *-2,07*	0,32	0,94 *0,62*	1,72	1,31 *0,25*	0,27 *0,60*	3,29 *0,65*	0,47 *0,99*	-7,62	0,08	53,3%	1996, M06 2002, M12	79	6	M

With *italics* we have denoted t-statistics with regard to estimates and respective probabilities with respect to test statistics as Jarque-Bera normality of residuals test (JB), Breusch-Godfrey serial correlation test (BG), conditional heteroscedasticity test (ARCH), White's test for heteroscedasticity (White), Chow stability test (Chow). The DW stands for Durbin-Watson test statistic, (A)DF for Dickey-Fuller unit root test, TP for turning points test statistic. The regression (3) was run with White's heteroscedasticity adjustment.

Source: own calculations.

Our preferred models of beta risk from economic and statistical point of view both for index WIG and $WIG20$ are those with interest rates and exchange rate PLN/US$. The results show that in the case of β_{WIG} within Sample I (see also equation 1 in Appendix) a moderate role was played by a difference of medium- and short-term interest rates. The difference of interest rates stands for risk premium and inflation expectations. As expected, the influence of interest rates was positive. A similar influence of interest rates we can notice in Sample II and in a full sample. We can, however, notice that the role of inflation expectations decreased with the advent of the period belonging to Sample II (see equations 2 and 3). In the end, we observed an impact of interest rates and not a term structure of interest rates. A different picture draws when looking at β_{WIG20} model estimates. The index $WIG20$ reflects price behavior of 20 biggest joint stock companies. We assume that prices and returns on stocks of those companies are more determined by fundamentals than by capital flows and speculation. The market for $WIG20$ is also more liquid. With respect to β_{WIG20} we observe an increasing role of inflation expectations in determining the beta risk (see also equations 4, 5 and 6). Results in Table 4 show that interest rate differential had a relatively strong impact on β_{WIG20} both in Sample II and in a full sample as measured by significance of respective estimates. We conclude that monetary policy, given its inflation-targeting behavior, and inflation expectations driven by this policy play a more important role with respect to $WIG20$ market than to WIG market. We link this with a more speculative behavior of smaller companies contained in the WIG index. As expected, an influence of exchange rate PLN/US$ turned out to be negative in both cases. This result is consistent with a literature discussed in previous sections. This effect should be attributed to a role of depreciation in improving the trade balance. It is evidenced for developed and emerging markets that in the short run it can lower the trade balance deficit. In the case of Poland, however, it is well documented (see e.g. Karadeloglou et al., 2001) that in a longer run devaluation feeds up inflation and the initial rise in competitiveness dies out rather quickly. Policy of devaluation should be in turn confronted with an important role of imports in the case of Poland, i.e. curbing imports by devaluation can be detrimental to the economy. Taking all this together we can conclude that contractory monetary policy in relation to the world could be an influential beta risk factor that increased the risk of domestic capital market in the analyzed period.

Based on statistics, we can see that our monetary models pass standard testing, i.e. we do not reject the normality of residuals, autocorrelation is not present, in most cases we do not detect ARCH effects and unconditional heteroscedasticity, and parameters are stable over time. We also have calculated ADF test statistics which show that residuals are stationary. Turning points statistics[8]

[8] The TP statistic is expressed as the number of matched by a model turns to tendency in a dependent variable to the number of all turns to its tendency.

(TP) are relatively high and denote that generally models follow ca. 50% of changes to tendency in dependent variables.

Now let us turn to an analysis of real factors that possibly influence the beta risk in Poland. We summarized the results in Table 5.

Table 5. Models of β_{WIG} and β_{WIG20} for real variables

Real model (R)																		
WIG																		
intercept	productivity	income	trade deficit	budget deficit	S_e	JB	DW	BG	ARCH	White	Chow	(A)DF	R^2 (adj.)	TP	sample	obs.	equation	forecasting model
-1,36	-0,03	x	0,10	0,03	0,33	1,37	1,73	0,51	0,61	6,60	0,66	-3,42	0,34	76,0%	1996, M07 – 1999, M01	31	7	
-2,96	*-1,72*	x	*3,83*	*1,79*		*0,50*		*0,47*	*0,43*	*0,68*	*0,77*							
-0,93	-0,05	-0,05	0,09	x	0,32	1,17	1,90	0,04	2,02	2,60	1,27	-5,19	0,39	60,0%		31	8	
-2,63	*-2,75*	*-2,69*	*3,82*	x		*0,56*		*0,84*	*0,15*	*0,86*	*0,33*							
-0,09	-0,01	x	0,02	0,01	0,31	0,22	1,51	4,66	0,03	15,37	0,76	-4,20	0,01	58,1%	1996, M07 – 2002, M12	78	9	R1
-0,29	*-1,40*	x	*1,29*	*0,69*		*0,90*		*0,03*	*0,86*	*0,08*	*0,80*							
0,07	-0,01	-0,01	0,01	x	0,31	0,27	1,54	4,26	0,12	10,54	0,64	-6,89	0,02	51,6%		78	10	R2
0,37	*-1,43*	*-1,25*	*1,09*	x		*0,88*		*0,04*	*0,73*	*0,31*	*0,91*							
WIG20																		
-0,94	-0,04	x	0,08	0,04	0,38	0,31	2,28	0,65	2,25	5,69	0,62	-2,89	0,25	69,6%	1996, M07 – 1999, M01	31	11	
-2,05	*-1,96*	x	*3,04*	*1,79*		*0,86*		*0,42*	*0,13*	*0,77*	*0,80*							
-0,81	-0,05	-0,06	0,08	x	0,34	0,25	2,00	0,02	1,43	7,59	0,76	-5,59	0,38	69,6%		31	12	
-2,16	*-2,80*	*-3,16*	*3,47*	x		*0,88*		*0,90*	*0,23*	*0,58*	*0,68*							
-0,16	-0,02	x	0,03	0,01	0,34	0,54	1,66	1,84	0,95	5,75	0,64	-7,67	0,06	50,8%	1996, M07 – 2002, M12	80	13	R1
-0,67	*-1,57*	x	*2,19*	*1,37*		*0,76*		*0,17*	*0,33*	*0,76*	*0,91*							
0,07	-0,01	-0,02	0,02	x	0,33	0,74	1,62	2,76	0,59	9,74	0,90	-7,24	0,05	48,3%		78	14	R2
0,32	*-1,48*	*-1,73*	*1,54*	x		*0,69*		*0,10*	*0,44*	*0,37*	*0,63*							

With *italics* we have denoted t-statistics with regard to estimates and respective probabilities with respect to test statistics as Jarque-Bera normality of residuals test (JB), Breusch-Godfrey serial correlation test (BG), conditional heteroscedasticity test (ARCH), White's test for heteroscedasticity (White), Chow stability test (Chow). The DW stands for Durbin-Watson test statistic, (A)DF for Dickey-Fuller unit root test, TP for turning points test statistic. The regressions: (9), (10) and (14) were run with White's heteroscedasticity adjustment.

Source: own calculations.

Again, we report our preferred model of beta risk both for index WIG and $WIG20$, i.e. models with the following risk factors: labor productivity, income, trade balance deficit, and budget deficit. Because of monthly data, income is proxied by industrial production since GDP is not reported on a monthly basis. We expressed the trade balance deficit and the budget deficit as relations to income. As we can easily see, productivity and income have a negative impact on the beta risk, both for β_{WIG} and β_{WIG20}. We conclude that a relative rise in competitiveness of the Polish economy may decrease country beta risk. We can also say that trade policies which do not put much emphasis on exports growth and expansionary fiscal policy are conducive to growth of risk. The deficits are traced by the market and their increase is perceived as endangering a stable economic growth. We could not find stable predictors for Sample II only. It is important to notice that models with real variables have in general higher ability to detect turns in tendency as evidenced by TP statistic which reach ca. 60%. Based on statistics, we see again that our real factor models pass standard testing, i.e. we do not reject the normality of residuals, autocorrelation is not present, in most cases we do not detect ARCH effects and unconditional heteroscedasticity, and parameters are stable over time. We also have calculated ADF test statistics which show that residuals are stationary.

Given our results, we obtained a puzzle. Both monetary and fiscal policies have direct and indirect impact on the pattern of risk of capital market. We should notice that exchange rates, prices, income, exports, imports which directly and indirectly influence beta risk are determined by the economic policy in Poland. This results in that we could hardly distinguish between purely monetary and real factors. A relative explanatory power of variables that we used in our analysis will be assessed in the procedure of checking predictive quality of econometric models. Section 3 is devoted to this problem.

3. Checking predictive quality of beta risk models

In this Section we will make an assessment of predictive quality of models with monetary and real factors analyzed in Section 2. We will follow a methodology proposed by Fair and Shiller (1990). Before applying a formal test, let us summarize *ex post* forecast errors for β_{WIG} and β_{WIG20} forecasts. Below in Table 6 we present *ex post* errors calculated for forecasts $\hat{\beta}_{WIG}$ obtained in a recursive procedure of one-period-ahead forecasting of beta risks based on preferred monetary (M) and real (R1 and R2) models. The out-of-sample testing period was 1999, M2 – 2002, M12.

Table 6. *Ex post* errors for $\hat{\beta}_{WIG}$

Index	Model	MAE	RMSE	MAPE	Theil	I_1^2	I_2^2	I_3^2	TP
	M	0,18	0,24	308,0%	0,42	5,1%	0,6%	94,4%	48,6%
WIG	R1	0,24	0,31	430,1%	0,48	18,0%	1,1%	80,9%	40,0%
	R2	0,23	0,30	362,3%	0,48	16,1%	0,9%	83,0%	40,0%

Source: own calculations.

As it can be seen, the model with monetary factors (M) had better forecasting quality than models with real factors (R1) and (R2). It is evidenced by favorable outcomes based on lower values of various measures of errors, lowest Theil's inequality coefficient. As for TP statistic, we obtained supportive results in the case of the monetary model. By comparison, within models with real factors we obtained better results in the case of model (R2). Below in Table 7 we present forecast errors for $\hat{\beta}_{WIG20}$.

Table 7. *Ex post* errors for $\hat{\beta}_{WIG20}$

Index	Model	MAE	RMSE	MAPE	Theil	I_1^2	I_2^2	I_3^2	TP
WIG20	M	0,23	0,30	326,8%	0,33	8,7%	0,0%	91,3%	48,6%
	R1	0,25	0,33	429,5%	0,38	9,9%	7,4%	82,7%	54,3%
	R2	0,26	0,35	384,3%	0,40	5,0%	1,2%	93,9%	34,3%

Source: own calculations.

As we can see, forecasts of β_{WIG20} generated by the model with monetary factors this time also turned out to be more accurate than forecasts generated by models with real factors. This accuracy is superior to the real factors models except for TP statistic as the model (R1) is best of all in matching changes to tendency in the risk variable. In general, however, real models perform worse than the monetary model.

For the purpose of quality assessment of forecasts generated by models of beta market risk we have applied a formal test proposed by Fair and Shiller (1990). Hence, we have estimated the following equation:

$$y_t - y_{t-1} = a_0 + a_1({}_{t-1}\hat{y}_{1t} - y_{t-1}) + a_2({}_{t-1}\hat{y}_{2t} - y_{t-1}) + u_t \qquad (2)$$

where ${}_{t-1}\hat{y}_{1t}$ denotes forecasts of y_t generated by the model 1, i.e. the model with monetary factors based on information available up to the moment $t-1$ with the use of recursive estimation for each period t. The predictor ${}_{t-1}\hat{y}_{2t}$ denotes forecasts generated accordingly by the model 2, i.e. the model with real factors, model (R1) or (R2) respectively, while u is an error term, $u \sim IN(0,\sigma_u^2)$. If neither model 1 nor model 2 contain any relevant information in terms of forecasts quality for variable y in period t, the estimates of a_1 and a_2 will be statistically insignificant. If both models generate forecasts that contain independent information, the estimates of a_1 and a_2 should both be statistically significant. If both models contain information but information contained in forecasts generated by model 2 is completely contained in forecasts generated by model 1 and furthermore model 1 contains additional relevant information, the estimate of a_1 will be statistically significant while the estimate of a_2 statistically insignificant. If both forecasts contain the same information, they are perfectly correlated and the estimation of parameters of (2) is not possible.

Now let us turn to applying a formal test of checking predictive quality of models (M), (R1) and (R2). Based on the models we have obtained in a recursive estimation one-period-ahead forecasts of beta risk. The forecasts are *quasi ex ante* forecasts as for the period t we have used all information available up to the period $t-1$. Furthermore, since forecasting models contain lagged explanatory variables, we did not have to forecast their values at time t to do *ex ante* forecasts of beta risk. In our analysis, as initial estimation sample, we have assumed the sample during 1996, M1-1999, M1. Then for the period 1999, M2-2002, M12 (47 observations) we have calculated one-period-ahead *quasi ex ante* forecasts based on forecasting models of beta risk adding one observation at a time and estimating the model after forecasts at time t were calculated. The forecasts were based on models denoted as (M), (R1) and (R2) (see Tables 4 and 5 and Appendix). Below in Table 8 we present estimation results of equation (2) for β_{WIG}.

Table 8. Estimation results of predictive quality model of β_{WIG} forecasts

WIG															
intercept	model M	model R1	model R2	S_e	JB	DW	BG	ARCH	White	Wald (M)	Wald (R)	R^2 (adj.)	TP	sample	obs.
-0,06	0,63	0,15	x	0,23	4,53	2,06	1,81	0,46	9,81	5,01	0,41	0,33	75,7%	1999, M02 2002, M12	47
-1,61	*2,24*	*0,64*	x		*0,10*		*0,18*	*0,50*	*0,08*	*0,03*	*0,52*				
-0,06	0,62	x	0,15	0,23	3,77	2,09	2,03	0,71	13,51	3,98	0,31	0,33	73,0%		
-1,57	*2,00*	x	*0,56*		*0,15*		*0,15*	*0,40*	*0,02*	*0,05*	*0,58*				

With *italics* we have denoted t-statistics with regard to estimates and respective probabilities with respect to test statistics as Jarque-Bera normality of residuals test (JB), Breusch-Godfrey serial correlation test (BG), conditional heteroscedasticity test (ARCH), White's test for heteroscedasticity (White), Wald coefficient restrictions test (Wald). The DW stands for Durbin-Watson test statistic, TP for turning points test statistic. The regressions were run with White's heteroscedasticity adjustment.

Source: own calculations.

The results show that generally forecasts generated by model (M) contain more relevant information than forecasts generated by models (R1) or (R2). We infer on the basis of t-statistics which say that coefficients by one-period-ahead *quasi ex ante* forecasts obtained in monetary models (M) are significant and coefficients by forecasts obtained in real models (R1) or (R2) are insignificant. The results are that information contained in forecasts by models (R1) or (R2) is completely contained in forecasts by model (M) and that model (M) contains additional information. We conclude that for beta risk β_{WIG} monetary factors as interest rates and exchange rate PLN/US$ were more influential than real factors as productivity, income, trade balance deficit and budget deficit as far as predictive quality of models is concerned. We have run Wald coefficient restrictions test assuming that a coefficient by model predictions equals to zero. The Wald test statistics say that we should reject the null in the case of model (M) and should not reject the null in the case of models (R1) or (R2). This says that only monetary factors influence the beta risk which makes forecasts more informative. The conclusions are consistent with those based on analysis of *ex post* errors.

Estimation results of equation (2) for beta risk β_{WIG20} are slightly different and we present them in Table 9.

Table 9. Estimation results of predictive quality model of β_{WIG20} forecasts

WIG20															
intercept	model M	model R1	model R2	S_e	JB	DW	BG	ARCH	White	Wald (M)	Wald (R)	R^2 (adj.)	TP	sample	obs.
-0,08	0,51	0,31	x	0,26	2,26	1,91	0,35	0,38	4,37	14,05	4,43	0,41	84,6%	1999, M02 2002, M12	47
-2,06	*3,75*	*2,11*	x		*0,32*		*0,55*	*0,54*	*0,50*	*0,00*	*0,04*				
-0,07	0,56	x	0,21	0,26	1,76	2,11	2,10	0,22	6,64	12,54	2,32	0,38	84,6%		
-1,76	*3,54*	x	*1,52*		*0,42*		*0,15*	*0,64*	*0,25*	*0,00*	*0,13*				

With *italics* we have denoted t-statistics with regard to estimates and respective probabilities with respect to test statistics as Jarque-Bera normality of residuals test (JB), Breusch-Godfrey serial correlation test (BG), conditional heteroscedasticity test (ARCH), White's test for heteroscedasticity (White), Wald coefficient restrictions test (Wald). The DW stands for Durbin-Watson test statistic, TP for turning points test statistic. The regression for (M) and (R2) was run with White's heteroscedasticity adjustment.

Source: own calculations.

As we can see *quasi ex ante* forecasts generated by model (M) and model (R1) both contain relevant information which is indicated by respective t-statistics. We should pay attention to that in the case of beta risk β_{WIG20}, i.e. the risk of the biggest companies, real factors as productivity, trade balance deficit and budget deficit, are more influential than in the case of beta risk β_{WIG}. It says that investors while making an assessment of the capital market performance in the case of the biggest companies are concerned with macro fundamentals which influence economic growth and hence influence stock prices. As we can see, on the base of Wald test, we should reject the null in the case of model (M) and model (R1) and should not reject the null in the case of model (R2). It says that factors as: interest rates, exchange rate, labor productivity, trade balance deficit and budget deficit excluding income, are factors that mostly influence the beta risk β_{WIG20}.

The statistical quality of equation (2) estimates both for β_{WIG} and β_{WIG20} is high. Generally, we can conclude that monetary variables as interest rates and exchange rates play a dominant role over real factors. The latter become, however, more and more influential in the case of country beta risk in Poland, especially in the market for big companies.

3. Conclusions

In this paper we have analyzed factors that possibly influence the market beta risk of Poland. We have estimated parameters of the market model of beta risk in which we have regressed returns on Polish stock market indexes WIG and $WIG20$ on world stock market indexes $DJIA$, $NASDAQ$, DAX and $FTSE$ on daily close-to-close data. The point estimates obtained in a daily sample within a month were averaged across all the models. Finally, we have obtained monthly time series of country beta risk measures in 84 observations. The beta risk variables were put as dependent variables in models of risk with monetary factors as interest rates and exchange rates and real factors as labor productivity, income, trade balance deficit and budget deficit as explanatory variables. Based on the monetary and real factors models and on analysis of *ex post* forecast errors and *ex ante* models of checking predictive quality, we conclude that in the case of beta risk β_{WIG} monetary variables were more influential than real variables in the period 1996, M1 – 2002, M12. As far as beta risk β_{WIG20} is concerned, we conclude that both monetary and real factors influenced the risk variable. This is to say that real factors are more influential in the case of the market for the biggest companies (index $WIG20$) than for all companies (index WIG) where short-term speculation plays more important role than analysis of market fundamentals.

We should point out that the integration of the Polish capital market with other European and world markets will be further strengthened by the accession of Poland into the EU. The accession itself should stabilize interest rates and exchange rates which is a pre-condition for adoption of the Euro currency. This is turn, given our results, should stabilize the capital market in Poland in terms of asset returns. Further studies should involve the structure of the Polish capital market, e.g. liquidity problems and the structure of capital involved, as well as impact of FDI and portfolio investments. An analysis based on sectoral stock indexes should give more insight into driving forces of the capital market in Poland.

The methodology applied and conclusions based on our analysis are consistent with studies in a large body of literature devoted to developed and emerging markets, i.e. we also managed to show that country beta risk of Poland is mostly influenced by financial variables as interest rates and exchange rates (see e.g. Erb at al., 1996; Groenewold and Fraser, 1997; Bracker and Koch, 1999; Gangemi at al., 2000; Goldberg and Veitch, 2002).

Appendix

Results of estimation

1 $$\hat{\beta}_{WIG,t} = 0{,}50 + 0{,}44[(i_{3m,t-3} - i^*_{3m,t-3}) - (i_{1m,t-3} - i^*_{1m,t-3})] - 0{,}10\Delta\log(S_{t-3})$$

2 $$\hat{\beta}_{WIG,t} = -0{,}05 + 0{,}02(i_{3m,t-5} - i^*_{3m,t-5}) - 0{,}01\Delta\log(S_{t-3})$$

3 $$\hat{\beta}_{WIG,t} = 0{,}05 + 0{,}02(i_{3m,t-1} - i^*_{3m,t-1}) - 0{,}03\Delta\log(S_{t-3})$$

4 $$\hat{\beta}_{WIG20,t} = -0{,}32 + 0{,}05(i_{3m,t-3} - i^*_{3m,t-3}) - 0{,}09\Delta\log(S_{t-3})$$

5 $$\hat{\beta}_{WIG20,t} = -0{,}16 + 0{,}05(i_{3m,t-4} - i^*_{3m,t-4}) - 0{,}02\Delta\log(S_{t-3})$$

6 $$\hat{\beta}_{WIG20,t} = 0{,}42 + 0{,}21[(i_{3m,t-5} - i^*_{3m,t-5}) - (i_{1m,t-5} - i^*_{1m,t-5})] - 0{,}03\Delta\log(S_{t-3})$$

7 $$\hat{\beta}_{WIG,t} = -1{,}36 - 0{,}03[\Delta\log(V_{t-5}) - \Delta\log(V^*_{t-5})] + 0{,}10tb_{t-3} + 0{,}03g_{t-2}$$

8 $$\hat{\beta}_{WIG,t} = -0{,}93 - 0{,}05[\Delta\log(V_{t-5}) - \Delta\log(V^*_{t-5})] - 0{,}05[\Delta\log(Y_{t-1}) - \Delta\log(Y^*_{t-1})] + 0{,}09tb_{t-4}$$

9 $\hat{\beta}_{WIG,t} = -0,09 - 0,01[\Delta\log(V_{t-5}) - \Delta\log(V^*_{t-5})] + 0,02tb_{t-3} + 0,01g_{t-2}$

10 $\hat{\beta}_{WIG,t} = 0,07 - 0,01[\Delta\log(V_{t-5}) - \Delta\log(V^*_{t-5})] - 0,01[\Delta\log(Y_{t-3}) - \Delta\log(Y^*_{t-3})] + 0,01tb_{t-4}$

11 $\hat{\beta}_{WIG20,t} = -0,94 - 0,04[\Delta\log(V_{t-5}) - \Delta\log(V^*_{t-5})] + 0,08tb_{t-4} + 0,04g_{t-3}$

12 $\hat{\beta}_{WIG20,t} = -0,81 - 0,05[\Delta\log(V_{t-5}) - \Delta\log(V^*_{t-5})] - 0,06[\Delta\log(Y_{t-1}) - \Delta\log(Y^*_{t-1})] + 0,08tb_{t-4}$

13 $\hat{\beta}_{WIG20,t} = -0,16 - 0,02[\Delta\log(V_{t-1}) - \Delta\log(V^*_{t-1})] + 0,03tb_{t-4} + 0,01g_{t-3}$

14 $\hat{\beta}_{WIG20,t} = 0,07 - 0,01[\Delta\log(V_{t-5}) - \Delta\log(V^*_{t-5})] - 0,02[\Delta\log(Y_{t-1}) - \Delta\log(Y^*_{t-1})] + 0,02tb_{t-4}$

Description of variables

β_{WIG}, β_{WIG20} – estimates of β parameters in a market model of country beta risk for stock market indexes WIG and $WIG20$,

i_{3m} – 3-month money market interest rate in Poland (%),

i_{1m} – 1-month money market interest rate in Poland (%),

i^*_{3m} – 3-month money market interest rate in U.S. (%),

i^*_{1m} – 1-month money market interest rate in U.S. (%),

S – exchange rate PLN/US$,

tb – relation of trade balance deficit to seasonally adjusted real industrial production in Poland (%),

g – relation of budget deficit to seasonally adjusted real industrial production in Poland (%),

V - labor productivity in Poland (%),

V^* - labor productivity in U.S. (%),

Y - seasonally adjusted real industrial production in Poland (%),

Y^* - seasonally adjusted real industrial production in U.S. (%).

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