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MARKET RISK DYNAMICS AND COMPETITIVENESS AFTER THE EURO:

Evidence from EMU Members

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MARKET RISK DYNAMICS AND COMPETITIVENESS AFTER THE EURO:

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Abstract

In this paper we propose an empirical model that considers theoretical facts on the relationship between real exchange rates and the net exports of the economy to supplement the interaction of a number of financial and economic factors with the stock market. We discuss the impact of exchange rate fluctuations on market risk in terms of Value at Risk (VaR). Our empirical findings show that common currency introduction produced increments in VaR whereas European stock returns are more sensitive to changes in competitiveness regarding the EMU rather than national exports. Finally, we show that the synchronisation of variation in competitiveness through the introduction of a single currency has made these changes more decisive in explaining financial market fluctuations.

KEYWORDS: Euro, Competitiveness, Market Risk, Net Export, Value-at-Risk, Volatility

JEL Codes: F33, G24, G28, O24

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1. Introduction

The financial landscape of the Europe has been growing rapidly over the last years. The introduction of the euro and the beginning of the single monetary policy on January 1, 1999, accelerated the pace of change. At the same time, the process towards European Economic and Monetary Union (EMU) has given a tremendous impetus to financial market integration in Europe. Capital controls were completely eliminated in the course of the 1980s and 1990s. The common currency introduction in 1999 removed all remaining exchange rate risk among the EMU participants, and marked the beginning of a single monetary policy for the euro area. It is widely believed that EMU will also greatly affect European capital markets via both direct and indirect effects (Danthine et al., 2001). Moreover, under assumption of common currency persistence as well as realization of certain conditions, the euro may well replace the dollar as international money (Dwyer and Lothian, 2002).

The degree of comovement among European equity markets seems to have increased as well. World equity markets have gone through various phases of integration. Returns in the world's major stock markets were highly correlated in the early 1900s, but then integration declined during the World Wars and in the 1970s (Goetzmann et al., 2001). However, since the 1980s and, in spite of the persistence in financial market segmentation (Guiso et al., 2004), global integration and the comovement in the world's major stock markets have steadily increased (Berben et al., 2005). Therefore, it is not surprising to see a link between the economic growth and financial markets coupled with the greater efficiency in the allocation of capital. More generally, such link is well seen in the functioning of markets, where the risk measures is of key importance (Hartmann et al., 2003 and Baele et al., 2004).

Besides the apparent rise in the degree of comovement among national financial markets, the European economy faced significant changes not only in financial markets but also in international competitiveness over the last decade. More recently, the growth of trading activity in financial markets coupled with numerous instances of financial instability and a number of widely publicized losses in financial institutions have resulted in a re-analysis of the risks. With ubiquitous risks and almost completely integrated financial markets in Europe, nowadays, one of the most important tasks of financial institutions is to evaluate the exposure to market risks, which is commonly done by estimating the Value at Risk.

Parallel with this development, turbulence in the foreign exchange markets has also undergone significant changes compared with the pre-euro period. This effect was foreseen by various economists (Ghironi and Giavazzi, 1997; Martin, 1997; Benassy et al., 1997; Gros and Thygessen, 1992; Kenen, 1995; Aglietta and Thygessen, 1995; Cohen, 1997).

But our main question is, were these two developments really correlated? And, if so, how exactly could monetary reform be held responsible for higher stock market risk? To answer this question we propose an empirical model that considers theoretical facts on the relationship between real exchange rates and the net exports of the economy to supplement the interaction of a number of financial and economic factors with the stock market. We also discuss the impact of exchange rate fluctuations on market risk in terms of Value at Risk.

The outline of the remaining sections will be as follows. In Section 2, the related literature is presented and the variables of the model are discussed. Section 3 presents market risk dynamics before and after the euro introduction. Section 4 presents our model describing the dynamics of stock market risk in competitiveness-exchange rates framework. Section 5 reports the empirical results and section 6 is the conclusion.

2. Related Literature

2.1. Exchange rates and stock markets

One can consider several potential links between exchange rates and stock market. For example, exchange rates may affect a firm's value by means of its impact on the liquidity of a firm's shares (Doukas et al., 2003; Huang and Stoll, 2001). In fact, there is a growing literature on the effect of liquidity on firm value. The pioneer work by Amihud and Mendleson (1986) present the first evidence to support the hypothesis that asset liquidity is priced in equilibrium. Among more recent papers, Datar et al. (1998), Brennan et al. (1998) and Easley et al. (1999) all suggest that asset liquidity affects a firm's value through its impact on the firm's expected return. Loderer and Roth (2005) strongly support the theory. Using data from the Swiss stock market and Nasdaq for a period 1995-2001, they estimate the effect of stock illiquidity, measured by the bid-ask spread, on stock prices. The authors employs Price Earning Ratio (PER), controlling for firm growth, dividend, risk and size finding that the larger is the bid-ask spread, the lover is the PER. Finally, Acharya and Pedersen (2005) employ the liquidity measure of Amihud (2002) and suggest a model which provides a unified framework for understanding the various channels through which liquidity risk may affect firms' value (see also Amihud and Mendleson (2006) for and ample surveys research on the effects of liquidity on asset prices and returns). In this context, if the asset liquidity, influenced by exchange rates, determines the firm's value and expected returns, then it is pertinent to study the link between the exchange rate and the market risk, which is the scope of this study.

However, the phenomenon of higher risk is not easily explained in such a straightforward context, as there is no obvious modification in this mechanism ascribable to the introduction of a currency. We consider stock prices and real exchange rates to be intermediated by changes in corporations competitiveness reflected in variations in trade flows directions. In turn, the changes in competitiveness are reflected in company's stock prices and related market risk.

In a multicountry world, movements in one exchange rate can be offset by other factors, such as movements in other exchange rates or interest rates. There are many studies that examine the relationship between exchange rate and international trade. For example, Asseery and Peel (1991) examine the influence of volatility on multilateral export volumes finding that exchange rates have significant positive effects on exports. At the same time, Bini-Smaghi (1991) finds strong support for the conventional assumption about volatility effects on trade. Cushman (1983), Kenen and Rodrick (1986), Giovannini (1988), Franke (1991), Pozo (1992), Sercu (1992), Sercu and Vanhulle (1992), Chowdhury (1993) and Kroner and Lastrapes (1993) among others, provide evidence that the level of exchange rate volatility impacts the volume of trade flows.

On the contrary, Koray and Lastrapes (1989), Lastrapes and Koray (1990), Gagnon (1993) in their studies on the effect of exchange rate volatility on trade conclude that the relationship between the volatility and trade is weak. Recently, Klaassen (2004) suggest to study the exchange rate effect on trade using data on countries with much more time variation in exchange rate risk, such as developing countries (as in Arize et al., 2000). Alternatively, the use of cross-sectional variation in exchange risk may be fruitful (De Grauwe and Verfaille, 1988). Campa et al. (2006) have performed an empirical analysis of transmission rates from exchange rate movements to import prices of the countries in EMU. They have estimated short and long-run elasticities for all euro countries, allowing them to change according to the type of product imported. The results obtained confirm that this transmission is high, although incomplete, in the short-run, and different across industries and countries. Long-run elasticities are higher, although estimated elasticities are still lower than unity, except for the traditionally more inflationary economies and for commodities. They conclude that, in general, the equality of pass-through elasticities among the different industries in each country or for the different countries given an industry cannot be rejected in the long-run. Moreover, it is accepted that if the volume of trade flow is impacted by exchange rate fluctuations so will the value of firms. But the conclusions of relevant empirical studies are quite different.

The estimation of the exchange-rate exposure began with the simple Jorion (1990) model evolved to more sophisticated time-varying models early this decade (e.g., Allayannis and Ihrig, 2001 or Bodnar et al., 2002). Amihud (1994) examines a sample of 32 top US exporters and concludes that their stock returns are not affected by changes in the value of the dollar. Bartov and Bodnar (1994) find that the abnormal returns of 208 firms are uncorrelated with changes in the value of the dollar. Griffin and Stulz (2001) noted that changes of weekly exchange rates had negligible impacts on industry stock indices in developed countries. In contrast, Bartov et al. (1996) finds that the return variability of US multinational corporations increases with an increase in exchange rate volatility. Bodnar and Gentry (1993), studying industry portfolios in the US, Japan and Canada, find that only 30% of them are significantly affected by exchange rate changes. He et al. (1996) examine a large sample of Japanese firms and find that of the 422 exporting companies, 25% are significantly affected by exchange rates fluctuations. Recently, Forbes (2002) documents how firms in 41 countries have their annual performance (measured as firm sales, net income, market capitalization and asset value) negatively affected over the span of exchange rate crises. Nevertheless, the discussions and arguments in the vast literature indicate that there is a relationship, which seems stronger or weaker in the light of different samples and studies.

In our opinion this interrelation between the exchange rate and corporation value is the one most likely to be the link between higher stock market risk and a common currency in the context of structural changes accounted after the euro, tested in the empirical section of our study.

Several potential factors of stock market risk are also included in our model in order to make it more specific. In particular, the remaining regressors (discussed in subsection 2.2) include proxies for business cycles, domestic market demand as well as bond yields, traded volume of stocks, and foreign reserves variables. Most of these factors are discussed in different contexts of interaction with financial market in financial and economic literature.

2.2. Variables discussion

The impact of different interest rates on stock returns is studied by a number of researchers (e.g. Gallant and Tauchen, 1997; Peiro, 1996) and a vast number of studies have treated the stock and bond markets in isolation. However, recently Kim et al. (2006) investigate stock and bond market integration over time within a common market jurisdiction motivated by the recent developments on stock–bond return co-movements in financial economics and the historical European Economic and Monetary union experience. Their study aims to examine whether the establishment of the EMU has induced a dynamic change in inter-stock–bond market integration by making inferences from the behavior of their daily conditional volatility interdependencies and time-varying conditional correlations. The authors conclude that as intra-stock and bond market integration with the EMU has strengthened in the sample period, interstock–bond market integration at the country level has trended downwards to zero and even negative mean levels in most European countries, Japan and the US. A similar study by Rapach et al. (2005), among other factors, reveals that relative long-term government bond yields have negative impact on real return from holding stocks. Pavlova and Rigobon (2003) identify interconnections between stock, bond and foreign exchange markets and characterize their joint dynamics as a three-factor model. Engsted and Tanggaard (2001) analyze the joint behaviour of Danish stock and bond markets over the period 1922–1996. They apply the same VAR methodologies as those developed by Shiller and Beltratti (1992) and Campbell and Ammer (1993) concluding that simple rational expectations present value models cannot explain the positive correlation between stock and bond returns apparently observed in the data.

Recent empirical work finds evidence of systematic movements in excess stock returns that are related to estimates of the underlying state of the business cycle (see Chauvet and Potter, 2000; Hamilton and Lin, 1996; Perez-Quiros and Timmermann, 1995; Whitelaw, 1994). The findings in these papers are that there is a strong relationship between stock market movements and business cycles. The stock markets usually begin the contractions some months before an economic recession and end before the trough, and, therefore, anticipate the economic recovery. That is, stock market fluctuations lead the business cycle and seem to be generated, to some extent, from expectations about changes in future economic activity. In this sense, Chauvet (1999) investigates the dynamic relationship between stock market fluctuations and the business cycle. The author concludes that the stock market is found to be a leading indicator of the state of the business cycle and can be used to anticipate turning points in real time.

Dumas et al. (2003) develop a ‘‘dynamic single-index’’ statistical model capturing the “world” business cycles as well as country-specific fluctuations. They consider current and past production as the information variable that investors use in their investment decision, as a way of predicting their decisions on which stage of the business cycle the economy is currently running. As stated in Boyd (2005) the unemployment news is a very important factor to explain the dynamics of financial markets. Similarly with Rapach et al. (2005) that also consider changes in the unemployment rate as a macroeconomic factor of stock returns, we use unemployment as a mirror of the business cycle[1] stage in our model.

The relation between trading volume and price pattern is among the more well-documented phenomena in financial research, which in turn suggests that volume could act as a suitable variable in the determination of the market risk. Cuñado et al. (2004) show that growth in traded volume, the next factor in our empirical model, has a significant impact on stock market volatility in Spain. They, however, conclude that it was not just the acceleration in trading volume that brought about the increased volatility but most likely the intensification of the process of economic development and opening the borders. Particularly, Conrad et al. (1994), Datar et al. (1998) and Brennan et al. (1998) have reported evidence of a negative relationship between volume and returns. In this sense, McMillan (2007) found substantial evidence of a negative relationship between volume and future returns, and that low volume is consistent with momentum behaviour in returns and high volume with reverting behaviour.

Ding et al. (2007) study the impact of traded volume in a different way. They investigate country-varying relation between trading volume and price pattern among short-horizon winners/losers in seven Pacific-Basin markets during the period 1990 to

[1] For a detailed discussion on the interdependence of unemployment rates and business cycles we refer to Bover et al. (2002) or Verho (2005).

2000. The results put forward the existence of monotonic relations between trading volume and short-horizon price pattern, which vary both among different countries and among winners/losers. These differences suggest that the relation between trading volume and price pattern need not be the same across countries, even among those in the same geographic region. Blume et al. (1994) argue that by studying volume data, which provides information on the quality or precision of information contained in past price movements, traders can find out useful information about the path of prices, such that lagged volume has a significant relationship with current returns. It is worth noting that link between the traded volume and volatility was found even in the studies which have made use of the high frequency data. For example, in their recent article, Darrat et al. (2007) investigate such data to test the dynamics of the causal relation between the traded volume and intraday volatility both in presence and absence of public news. Indeed, they found bi-directional causality between trade volume and return volatility in the period with public news. In the absence of public news, they argue that the causality flows positively and significantly from volume to volatility without feedback.

As pointed out in Palley (2002) the domestic demand, another explaining variable used in our model, rests on four pillars: (1) improved income distribution, (2) good governance, (3) financial stability and space for counter-cyclical stabilization policy, and (4) an adequate fairly priced supply of development finance. Gürkaynak and Wolfers (2005) argue that retail trade and business confidence level have predictive power in financial markets. Thus, given the importance of the domestic demand and to reflect the process of economies development, a proxy variable (changes in retail trade) is considered.

In theory, the volume of international financial transactions, and therefore reserve holdings, should increase with economic size. At the same time, the determinants of foreign reserve holdings can be grouped into five categories which can be summarized as: economic size, capital account vulnerability, current account vulnerability, opportunity cost and finally exchange rate. Potentially, the explanatory variables for each on those categories may be, GDP to reflect the economic size, ratio of capital flows or broad money to GDP, short-term external debt, foreigners' equity position to cover the capital account vulnerability, the share of imports or exports in output, volatility of export receipts as a explanatory variable of current capital vulnerability, interest rate differentials as opportunity cost and exchange rates dynamics

(Gosselin, M-A., Parent, N., 2005). As a result, the holdings of foreign reserves may directly impact on the competitiveness of a country.

Empirical research on international reserves establishes a relatively stable long-run demand for reserves (Lane and Burke, 2001). At the same time, it is argued that an ample part of the foreign exchange reserves is usually invested in international financial markets (mainly in the liquid bond markets) and consistently the changes in the volumes of reserves will somehow be reflected in the financial market volatility. For example, Mendoza (2004) concludes that reserve management is motivated by a desire to self-insure against a financial crisis. Thus, covering this variable which potentially may impact on general stability of the currency and financial markets (Masson and Turtleboom, 1997; Lehay, 1996; Hening, 1997) is also considered in our study.

Therefore, given the importance of the factors mentioned above in the competitiveness-exchange rates framework, our research discusses how the latter explain the market risk dynamics in a sample of EMU countries. The empirical results make it possible to obtain additional findings on how the competitiveness of companies and stock markets interact within the sample of the countries under consideration.

3. Market risk dynamics in pre- and post-euro periods

Financial risk is the prospect of financial loss (or gain) due to unforeseen changes in underlying factors. The changes that euro introduction in 1999 caused in stock markets is the target of particular study. To evaluate the market risk before and after the euro we used the Value at Risk indicator (see e.g. Jorion, 2000; Goorbergh and Vlaar, 1999). Value at Risk (VaR) is defined as the maximum potential change in value of a portfolio of financial instruments with a given probability over a certain time horizon, with the assumption that the composition of the theoretical portfolio remains the same[2]. VaR measures have many applications, such risk management and for regulatory requirement and it can be utilized as a vehicle for corporate self-insurance since VaR can be interpreted as the amount of uninsured loss acceptable to a

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[2] Analytically, the VaR is defined by the top limit of integral of the probability density function (P) of expected returns (r) $\alpha = \int_{-\infty}^{E(r)-VaR} P(r)dr$. Usually it is assumed that the expected value of the returns is zero so we can transform $\alpha = \int_{-\infty}^{E(r)-VaR} P(r)dr$ into $\alpha = \int_{-\infty}^{-VaR} P(r)dr$

corporation[3] (Shimko, 1997). In this framework, the Basel Committee on Banking Supervision (1996) requires financial institutions such as banks and investment firms to meet capital requirements based on VaR estimates.

Estimating volatility is the essence of evaluating of market risk. Among the variance methods of VaR estimation the static models do not take volatility clustering into account. In this sense, Piñeiro et al. (2006) have shown how the non-static models are more suitable for the variance prediction. By far the most popular model which captures this phenomenon is the Generalized Autoregressive Conditional Heteroskedasticity (GARCH), introduced by Bollerslev (1986) as an extension of the Autoregressive Conditional Heteroskedasticity (ARCH) model by Engle (1982). The GARCH model defines an innovation η_{t+1}, i.e., some random variable with mean zero conditional on time t information, I_t. This time t information is a set including not only the innovation at time t, $\eta_t \in I_t$, and all previous innovations, but also any other variable available at time t as well. In finance theory, η_{t+1} might be the innovation in a portfolio return. In order to capture serial correlation of volatility, or volatility clustering, the GARCH model assumes that the conditional variance of the innovations depends on the latest past squared innovations as is the assumption in the less general ARCH model, possibly augmented by the previous conditional variances. In its most general form, GARCH(p,q), can be written as:

$$\sigma_t^2 = \omega + \sum_{j=1}^{p} \beta_j \sigma_{t-j}^2 + \sum_{i=1}^{q} \alpha_i \eta_{t-i+1}^2 \qquad (1)$$

p lags are included in the conditional variance, and q lags are included in the squared innovations. In our study we regard these innovations as deviations from some constant mean portfolio return:

$$r_{t+1} = \mu + \eta_{t+1} \qquad (2)$$

expressed η_{t+1} as $\sigma_t \varepsilon_{t+1}$, where ε_{t+1} is assumed to follow some probability distribution with zero mean and unit variance, such as the standard normal distribution. The

[3] For example, a corporation should buy external insurance when the self-insurance losses, reflected by VaR measures, are greater than the cost of insurance by hedging.

parameters are conditioned as $\omega > 0$, $\beta \geq 0$ and $\alpha \geq 0$ to ensure positive variances. If the market was volatile in the current period, the next period's variance will be high, and is intensified or offset in accordance with the magnitude of the return deviation this period. Naturally, the impact of these effects hinges on the parameter values. Note that for $\alpha + \beta < 1$, the conditional variance exhibits mean reversion, i.e., after a shock it will eventually return to its unconditional mean $\omega/(1-\alpha-\beta)$. In this way, if $\alpha + \beta = 1$, this is not the case, we would have persistence.

In order to estimate these parameters by means of likelihood maximisation, one has to make assumptions about the probability distribution of the portfolio return innovations η_{t+1}.

We consider Gaussian innovations

$$\varepsilon_t \overset{iid}{\sim} N(0,1), \qquad \eta_{t+1}|I_t \sim N(0,\sigma_t^2) \tag{3}$$

leading to a conditional log likelihood of η_{t+1} equal to:

$$\ell_t(\eta_{t+1}) = -\log\sqrt{2\pi} - \frac{1}{2}\log\sigma_t^2 - \frac{\eta_{t+1}^2}{2\sigma_t^2} \tag{4}$$

The log-likelihood for all series is $\sum_{t=1}^{T} \ell_t(\eta_{t+1})$.

The GARCH(1,1) was used to predict the volatility dynamics during VaR estimation period for a sample of ten EMU member states, since it is found to be adequate for many financial time series (Bollerslev, Chou and Kroner, 1992). McNeil and Frey (2000) use GARCH in yet another way to get value at risk. They use GARCH to estimate the volatility, and extreme value theory to get tail probabilities. Ahlstedt (1998) argues that the GARCH models represent a methodological and empirical improvement over other estimates. Therefore, the estimated impact of changes in Euro/USD exchange rates on net exports of EMU countries to the USA is the key regressor of our interest explaining the dynamics of the level of market risk in our empirical model.

The daily VaR estimates, for left tail probability of 1% according to Basel Accord (1996) are reflected in figure 1 in appendix 1 while the average VaR for the pre- and post- euro periods and the corresponding growth in absolute terms is reported in the table 1. The increase in average daily VaR is obvious in EMU major stock markets. Among the countries with significant growth in market risk are the two largest economies of the EMU – Germany and France, only Italy and Austria produced a slight reduction in VaR.

In the context of our study, the exchange rates dynamics is of high importance because it affects decisions of market participants. The consequences of exchange rate volatility on trade have long been at the center of the debate, and its impact on stock markets was widely studied. In fact, our calculations confirm a significant increase in the volatility of the USD/EURO real exchange rate between 01/1995 and 08/2004 period. Figure 2 shows the volatility of exchange rate measured as a 5-month annualized standard deviation of the first differences of the log real exchange rates (USD/EUR). Similar to Bagella et al. (2004), we are interested to show the volatility dynamics, and not in the investigation of its law of variation. Therefore, we prefer the historical volatility calculation rather ARCH or GARCH measures. As is seen from the linear fit (see Fig. 2), the growth tendency has been maintained over the whole considered sample. The maximum degree of volatility has been reached on 05/2000 with 0.1357 points while the minimum has been on 04/2001 with 0.0161 points. Here, we also report the logarithmic fit (see Fig. 2) which confirms the persistence in the volatility growth[4].

Table 1
VaR before and after euro and the growth in absolute terms

Country	**Index**	**Exante (%)** (1995/01-1998/12)	**Expost (%)** (1999/01-2004/08)	**Growth** (% points)
Germany	DAX30	-2.97	-3.97	1.00
Belgium	BEL20	-2.16	-2.76	0.60
France	CAC40	-2.94	-3.50	0.56
Ireland	ISEQ40	-2.09	-2.55	0.46
Spain	IBEX35	-2.96	-3.36	0.40
Finland	HEX25	-3.53	-3.88	0.35
Portugal	PSI20	-2.31	-2.45	0.14
Netherlands	AEX24	-2.66	-2.78	0.12
Italy	MIB30	-3.43	-3.19	-0.24
Austria	ATX20	-2.42	-2.18	-0.24

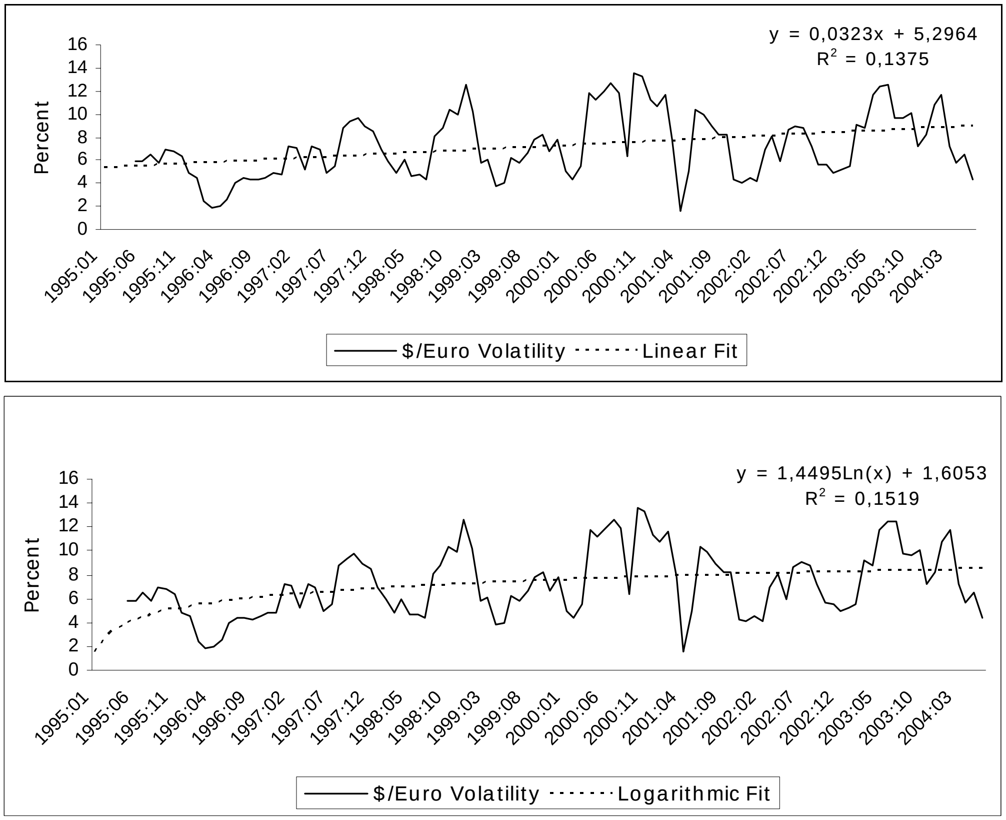
Note: For normal distribution assumption of returns VaR is computed as: $VaR = -V\left(e^{\mu+\sigma\phi^{-1}(\alpha)} - 1\right)$, where V represents the initial value of some theoretical portfolio and $\phi(\cdot)$ is the cumulative distribution function of the standard normal probability distribution. μ and σ with GARCH(1.1) are the estimates of the parameters of normal probability distribution function.

Source: Our own estimates based on Reuters data.

[4] We obtained the same result from the second order polynomial fit but do not report it to avoid the overloading of the paper.

Further, we construct and apply an empirical model to explain how the euro introduction could impact the stock market risk.

Figure 2: \$/Euro annualized volatility



Note: Real exchange rates are derived by multiplying the nominal exchange rate by the ratio of the U.S. to local currency CPI. Index (2000 average=100%). The volatility was calculated by a rolling standard deviation with a 5 month moving window.
Source: Authors' own calculation based on monthly series from Financial Statistics of the Federal Reserve Board and International Financial Statistics of the IMF.

4. Empirical Model

The starting point is the relationship between financial market risk (ϕ), estimated on stock price volatility, and a sample of explaining variables – changes in exchange rates (ε), changes in domestic market demand (λ), traded volume of stocks (ν), bond yields (τ), foreign official reserves (ϖ) and the business cycles (ρ).

$$\phi = \phi\left(\Delta\varepsilon, \Delta\lambda, \nu, \tau, \varpi, \rho\right) \tag{5}$$

We assume that the main link between the stock market risk and exchange rates, which may be affected by the common currency introduction, is the change in general competitiveness of the economy, reflected in terms of changes in net exports.

The relationship between real exchange rates and net exports is widely discussed in the financial literature. A number of comparatively older studies (e.g. Ethier, 1973; Cushman, 1986; Peree and Steinherr, 1989) have shown that an increase in exchange rate volatility will have adverse effects on the volume of international trade. More recent studies have demonstrated that increased volatility can have ambiguous or positive effects on trade volume (Viaene and de Vries, 1992; Franke, 1991; Sercu and Vanhulle, 1992). Barkoulas et al. (2002) concludes that under risk aversion, the benefits of international trade are reduced, resulting in a decrease in the volume of international trade. The trade surplus or deficit is reduced as well. However, they note that analysis which considers only the (often indeterminate) effects of exchange rate uncertainty on the volume of trade will not be capable of generating predictions of optimal behaviour.

Our interest in this relationship is limited to the most general ideas on the interaction of net exports with the exchange rates dynamics by estimating the impact of changes on net export, without any requirement of model modifications or prediction making.

Relating the macroeconomic dependence of import (φ) and export (ι) with the exchange rates, GDP (ψ) and GDP of the counterpart (ψ') we have:

$$\xi = (\varphi - \iota) = \varphi\left(\overset{-}{\varepsilon}, \overset{+}{\psi'}\right) - \iota\left(\overset{+}{\varepsilon}, \overset{+}{\psi}\right) = \xi\left(\overset{-}{\varepsilon}, \overset{-}{\psi}, \overset{+}{\psi'}\right) \tag{6}$$

Hence, the net export (ξ) changes caused by the exchange rate fluctuations from equation 6 could be expressed as $\left(\Delta\varepsilon \quad \left({}^{\partial\xi}\!/\!{}_{\partial\varepsilon}\right)\right)$.

Thus, our model describing the dependence of market risk from factors including changes in competitiveness for a single country is:

$$\phi = a_0 + a_1\left(\frac{\partial \xi}{\partial \varepsilon}\right)\Delta\varepsilon + a_2\Delta\lambda + a_3\nu + a_4\tau + a_5\varpi + a_6\rho \qquad (7)$$

These particular changes in net exports reflect the changes in competitiveness of the output of the country vs. the output of the trade party. Hence, the proxy for the general competitiveness of EMU countries is the change in the EMU net exports ($\hat{\xi}$) equal to:

$$\Delta\hat{\xi} = \sum_{i=1}^{n}\left(\Delta\varepsilon_i \frac{\partial \xi_i}{\partial \varepsilon_i}\right) \qquad (8)$$

The main assumption is that after euro introduction the changes in net exports of all the member states reflect the fluctuations of the single currency $(\hat{\varepsilon})$.

$$\Delta\hat{\xi} = \Delta\hat{\varepsilon}\sum_{i=1}^{n}\frac{\partial \xi_i}{\partial \hat{\varepsilon}} \qquad (9)$$

Thus, the changes in net exports of separate countries caused by the exchange rate changes are of the same sign. A single currency has a synchronising effect on general competitiveness changes, so that EMU has a larger $\Delta\hat{\xi}$ in the case of the euro. By replacing this term in the equation (7) for the $i-th$ term from the n countries we obtain:

$$\phi_i = a_0 + a_1\left(\Delta\hat{\varepsilon}\sum_{i=1}^{n}\frac{\partial \xi_i}{\partial \hat{\varepsilon}}\right) + a_2\Delta\lambda_i + a_3\nu_i + a_4\tau_i + a_5\varpi_i + a_6\rho_i \qquad (10)$$

From that our proposition is that the exchange rate driven changes of general competitiveness determine the level of financial market risk, which explains the phenomenon of higher value-at-risk in case of a vulnerable euro. These ideas are summarized following two propositions.

Proposition I.

In case of a single currency the $\sum_{i=1}^{n}\left(\Delta\varepsilon_i \frac{\partial\xi_i}{\partial\varepsilon_i}\right)$ variable is replaced with $\Delta\hat{\varepsilon}\sum_{i=1}^{n}\frac{\partial\xi_i}{\partial\hat{\varepsilon}}$, where $\left|\Delta\hat{\varepsilon}\sum_{i=1}^{n}\frac{\partial\xi_i}{\partial\hat{\varepsilon}}\right| \geq \left|\sum_{i=1}^{n}\left(\Delta\varepsilon_i \frac{\partial\xi_i}{\partial\varepsilon_i}\right)\right|$ because of the synchronized impact on foreign trade. The currency fluctuations cause greater fluctuation in general competitiveness of EMU production and result in higher volatility and risk in stock markets.

Proposition II.

The more significant is the variable $\sum_{i=1}^{n}\left(\Delta\varepsilon_i \frac{\partial\xi_i}{\partial\varepsilon_i}\right)$ (compared with $\Delta\varepsilon_{it}\left(\partial\xi_i/\partial\varepsilon_i\right)$ national alternative) in $\phi_i = a_0 + a_1\left(\sum_{i=1}^{n}\left(\Delta\varepsilon_i \frac{\partial\xi_i}{\partial\varepsilon_i}\right)\right) + a_2\Delta\lambda_i + a_3\nu_i + a_4\tau_i + a_5\varpi_i + a_6\rho_i$ equation, the deeper are particular economies integrated, and thus the euro fluctuations are more decisive for particular stock markets.

To test proposition I empirically, it is sufficient to prove the significance of the ε in the eq.6. Therefore, when the empirical results support proposition II, together with higher volatility of real exchange rates in the post-euro period, we can fully explain the indicated growth in VaR after the euro.

5. Empirical Findings

5.1. Changes in competitiveness vs. exchange rates

Before proceeding to the empirical testing of the stated hypothesis explaining the dynamics in the level of market risk we need to obtain estimated changes in net export. We used balanced monthly panel data 1995/01-2004/06 (see table 5 in appendix 2) for 11 EMU member countries to build an empirical model where the counterpart of the EMU is the USA. In context of our study the appropriate panel regression model has fixed individual effects (b_{i0}) and different slopes (Cornwell and Schmidt, 1984) for log-exchange rates.

$$\xi_{it} = b_{i0} + b_{i1}\ln\varepsilon_{i(t-l)} + b_2\left(\frac{\Psi'}{\Psi_i}\right)_{t-l} \tag{11}$$

Heteroskedasticity adjusted estimates of the model are reported in Table 2.

Prior to proceeding with our empirical findings, we perform Mancini-Grifoli and Pauwels (2006) procedure in order to detect a possible structural change in the European competitiveness after the euro introduction in 1999. This new procedure offers three main practical and technical advantages over others. First, the test does not make any distributional assumptions as it estimates empirically the distribution of the test statistic using an empirical subsampling methodology. Second, the power of the test remains high even when there are very few observations after the break date. Third, the test requires very few regularity conditions. It remains asymptotically valid despite non-normal, heteroskedastic and/or autocorrelated errors, and non strictly exogenous regressors.

Hence, the regression that serves as the basis for test of structural break has the following general form:

$$Y_{it} = \begin{cases} X_{it}'\beta_0 + U_{it} & t = 1,\ldots,T \\ X_{it}'\beta_{1t} + U_{it} & t = T+1,\ldots,T+m \end{cases} \qquad (12)$$

for individuals $i = 1,\ldots,n$ and where T is the supposed break date. The test hinges the next hypotheses: $H_0 : \beta_{1t} = \beta_0$ against $H_A : \beta_{1t} \neq \beta_0$. In order to build the test the authors consider more observations after the break date than regressors d, $(m \times n) \geq d$. Briefly, the test statistic is a positive definite quadratic form obtained from the transformed $(m \times n) \geq 1$ vector of residuals by the $(m \times n) \times (m \times n)$ covariance matrix, projected onto the column space of $(m \times n) \times d$ matrix of transformed post-instability regressors. As authors argue, the equivalent of the generic test statistic in Andrews (2003) for panel data can be defined after considering an interval τ_r which goes from $[r, r+m-1]$ and where $r \in \{1,\ldots,T+1\}$, as:

$$S_r(\beta,\Sigma) = A_r(\beta,\Sigma)' V_r^{-1} A_r(\beta,\Sigma), \qquad (13)$$

$$A_r(\beta,\Sigma) = X_{\tau r}' \hat{\Sigma}_{T+m}^{-1} W_{\tau r}, \qquad (14)$$

$$V_r(\Sigma) = X_{\tau r}' \hat{\Sigma}_{T+m}^{-1} X_{\tau r} \qquad (15)$$

with $\hat{W}_{\tau_r} = \left(\mathrm{Y}_{\tau_r} - \mathrm{X}_{\tau_r}\beta\right)$ where $\hat{W}_{\tau_r}$ is the $(m \times n) \times 1$ residual vector of observations starting at r, with $\beta = \hat{\beta}_{T+m}$ defined to be the coefficient vector estimated over the $T+m$. The variance-covariance matrix, $\hat{\Sigma}_{T+m}$, is given by,

$$\hat{\Sigma}_{T+m} = (T+1)^{-1} \sum_{r=1}^{T+1} \left(\hat{U}_{\tau r} \hat{U}'_{\tau r}\right) \tag{16}$$

where the $(m \times n) \times 1$ residual vector, $\hat{U}_{\tau_r}$, is defined as $\hat{U}_{\tau_r} = \left(\mathrm{Y}_{\tau_r} - \mathrm{X}_{\tau_r}\hat{\beta}_{T+m}\right)$. Noteworthy that this covariance matrix corrects for serially correlated errors, hetereskodasticity and potential cross-sectional correlation.

The particular form of the test statistic for the post-break residuals is defined as:

$$S = S_{T+1}\left(\hat{\beta}_{T+m}, \hat{\Sigma}_{T+m}\right) \tag{17}$$

At the same time, the critical values, S_r, are found by empirically generating a distribution function for the statistic under the null of stability. As before, if $(m \times n) \geq d$ the $T - m + 1$ different S_r values are defined as:

$$S_r = S_r\left(\hat{\beta}_{2,(r)}, \hat{\Sigma}_{T+m}\right) \tag{18}$$

where $\hat{\beta}_{2,(r)}$ is estimate the of β over $t = 1,...,T$ observations but excluding $\frac{m}{2}$ observations. The optimization of the power and size is the reason behind such exclusion, compared with the exclusion of only m observations or no observations at all.

However, the variance-covariance matrix , $\hat{\Sigma}_{T+m}$, as defined above will not be invertible in most cases, as it will in general not be of full rank, and thus for its adaptation to the panel data requires certain restrictions on the $(m \times n) \times (m \times n)$ covariance matrix to make in invertible. Hence the covariance matrix is redefined assuming sectional interdependence although continue to allow for serial correlation and heteroskedasticity. The redefined matrix has the following expression:

$$\widetilde{\Sigma}_{T+m} = (T+1)^{-1} \sum_{r=1}^{T+1} \left(\hat{U}_{\tau_r} \hat{U}'_{\tau_r} \right) = \hat{\Sigma}_{T+m} \tag{19}$$

except that $E\left[U_{i},_{\tau_r} U'_{j},_{\tau_r} \mid \mathrm{X}_{ij}\right] = 0$, for $i \neq j$ with $i, j = 1,\ldots,n$ and $U_{i},_{\tau_r}$ is an $m \times 1$ vector made up of the elements in $U_{\tau r}$ corresponding to individual i. The resulting covariance matrix $\widetilde{\Sigma}_{T+m}$ is block diagonal. Each block corresponds to an individual in the panel, and it is thus of dimension $(m \times m)$. Since the inverse of a block diagonal matrix is the inverse of each of its blocks, the condition for invertibility is satisfied[5].

Among other tests, an important advantage of this one is that it does not require normal iid errors and strictly exogenous regressors, while the F-type tests do.

The reported statistics (see table 2) suggests that the euro introduction cause a structural break in the European competitiveness.

Based on the b_{i1} vector from eq. 11 and the log-returns of the exchange rates with the five month lag, the impact of the exchange rate fluctuations on the net export of the particular countries (the $\Delta\varepsilon_{it}$ $\left({}^{\partial \xi_i}\!\big/_{\partial \varepsilon_i} \right)$ series) is estimated. We interpret these estimates as changes of competitiveness of domestic production in the international market (considering US market). Finland and Ireland are removed from the sample of the countries during further analysis because of insufficient observation during the period of study. At the same time because of non robust b_{i1} coefficient, the Luxembourg is also excluded from the group.

It is normal to assume that the larger the $\Delta\hat{\xi}_t$ caused by FX changes, the stronger is the position of European companies' shares at the stock markets. Therefore investors can expect the related market risk (VaR) to fall.

[5] See Mancini-Griffoli and Pauwels (2006) for detailed computations of alternative conditions for the inversion of the covariance matrix.

Table 2

FGLS estimates of the model (eq.11)

Dependent Variable: ξ_{it}

Country (i)	b_{i0}	b_{i1}	b_2
Common			0.274* (2.334)
Country Specific			
Austria	883.791	-172.721** (-2.860)	
Belgium	1594.762	-422.875** (-5.282)	
Finland	1391.424	-278.212** (-7.341)	
France	6368.738	-1219.168** (-7.106)	
Germany	16010.822	-2919.492** (-6.719)	
Ireland	11648.354	-2339.249** (-7.451)	
Italy	5265.262	-898.207** (-6.374)	
Luxembourg	-421.855	56.980 (1.140)	
Netherlands	147.510	-222.84* (-1.976)	
Portugal	384.072	-78.044** (-2.727)	
Spain	808.284	-181.321** (-2.675)	
l (*lag*)		5	6
Unweighted Statistics			
Adj. R-sq.	0.881	S.E. of regression	285.020
Significance of Group Effects Test			
F-stat	34.605 [a]	F-crit. (1%)	2.336
White General Test			
Chi-sq. stat	22.834 [b]	Chi-sq. crit (1%)	15.086
Mancini-Griffoli and Pauwels test[c]			
S-stat	73.08[c]	S_r (1%)	72.84
Included Observations			
Total panel obs.	1188	Obs. in cross sections	108

Note:

a) $H_0 : b_{11} = ... = b_{n1}$ of common constant term is rejected. We use the regression model with fixed individual effects as all the results are to be applied only on a sample of EMU countries.

b) H_0 of homoskedasticity is rejected.

c) Mancini-Grifolli and Pauwels test of structural break for panel data. The null hypothesis that there is no structural change over the period 1995/01-2004/06 is rejected.

t-stats. are given in the parentheses.

** significant at 1%, * significant at 5% confidence level.

5.2. Explaining higher stock market risk

5.2.1. The choice between two parallel models

Certain proxies are used for the variables in eq. 10 along with estimated proxy of changes of general $\left(\Delta\hat{\varepsilon}_{(t-5-l)}\sum_{i=1}^{n} b_{i1}\right)$ and alternatively country individual $\left(\Delta\varepsilon_{i(t-5-l)}b_{i1}\right)$ competitiveness because of real exchange rate fluctuations. The changes in retail trade volumes are used to proxy the dynamics of domestic market demand ($\Delta\lambda$). We also use the long-term government bond yields, the importance of which already has been discussed (γ). Unemployment rate is included to reflect the particular stage of business cycle (ρ). The reason behind this is that the higher is the unemployment, the deeper is the crisis and higher is the market risk.

$$\phi_i = a_0 + a_1\left(\Delta\hat{\varepsilon}_{(t-5-l)}b_{i1}\right) + a_2\Delta\lambda_{i(t-l)} + a_3 \ln(\nu_{i(t-l)}) + a_4\gamma_{i(t-l)} + \\ + a_5 \ln(\varpi_{i(t-l)}) + a_6\rho_{i(t-l)} \qquad (20)$$

$$\phi_i = a_0 + a_1\left(\Delta\hat{\varepsilon}_{(t-5-l)}\sum_{i=1}^{n} b_{i1}\right) + a_2\Delta\lambda_{i(t-l)} + a_3 \ln(\nu_{i(t-l)}) + a_4\gamma_{i(t-l)} + \\ + a_5 \ln(\varpi_{i(t-l)}) + a_6\rho_{i(t-l)} \qquad (21)$$

We consider two identical models by taking the country individual competitiveness variable in one (eq. 20) and the general competitiveness in the other (eq. 21) case (see table 3). Balanced monthly panel data for post euro period (1999/01-2003/12) has been used[6] (see table 6 in appendix 2). The results suggest that replacing the $\Delta\hat{\varepsilon}_{(t-5-l)}b_{i1}$ in the first model with the $\Delta\hat{\varepsilon}_{(t-5-l)}\sum_{i=1}^{n} b_{i1}$ in the second one improves the model. If the first variable is significant at a 5% confidence level, the variable of general competitiveness is significant at a level of 1%. We also perform a model considering both individual and general competitiveness variables jointly (see eq. 22) in order to corroborate our findings from the two parallel models (see eq. 20 and 21).

[6] Last six months were dropped due to the balanced data use.

$$\phi_i = a_1\left(\Delta\hat{\varepsilon}_{(t-5-l)} b_{i1}\right) + a_2\left(\Delta\hat{\varepsilon}_{(t-5-l)} \sum_{i=1}^{n} b_{i1}\right) + a_3\Delta\lambda_{i(t-l)} + a_4 \ln(\nu_{i(t-l)}) + a_5\gamma_{i(t-l)} + $$
$$+ a_6 \ln(\varpi_{i(t-l)}) + a_7\rho_{i(t-l)} \qquad (22)$$

It is confirmed, based on numbers presented in table 4, how the general competitiveness still remains statistically more significant than the individual competitiveness variable.

Thus the empirical results show that the growth in real exchange rates reduces the international competitiveness of particular economies exports, and vice versa, according to the macroeconomic theory.

We show that the changes in competitiveness in turn cause fluctuations in the level of stock market risk by increasing the risk when the national production loses position on the international markets, and by calming down the stock market when competitiveness grows.

Table 3.
FGLS Estimates of alternative models (eq.20 and eq.21)

Dependent Variable: ϕ_i

Model	(1)	(2)	l (lag)
Constant term	-1.601 (-0.837)	-1.841 (-0.968)	
Competitiveness change	-1.91E-03* (-2.311)	-2.65E-04** (-2.647)	0
Change in domestic demand	0.016** (2.732)	0.016** (2.731)	3
Traded stock volume [c]	0.132* (1.953)	0.128 (1.894)	1
Bond yields	-0.396** (-4.564)	-0.402** (-4.623)	0
Foreign reserves [c]	0.342 (1.532)	0.375 (1.684)	1
Unemployement	0.144** (2.753)	0.143** (2.717)	0
AR(1)	0.746** (24.399)	0.749** (24.621)	
Unweighted Statistics			
Adj. R-sq.	0.603	0.604	
S.E. of Regression	0.842	0.841	

Significance of Group Effects Test		
F-stat	1.1424[a]	1.1276[a]
F-crit. (1%)	2.6772	2.6772
White General Test		
Chi-sq. stat	29.6992[b]	28.1000[b]
Chi-sq. crit (1%)	27.6882	27.6882
Included Observations		
Total panel obs.	480	480
Obs. in cross sections	61	61

Note:
a) $H_0 : b_{11} = ... = b_{n1}$ of common constant term is accepted.
b) H_0 of homoskedasticity is rejected.
c) Variables are expressed in logs.
t-stats. are given in the parentheses.
** significant at 1%, * significant at 5% confidence level.

Hence, the growth in exchange rates results in higher stock market risk. A set of other factors of stock market risk and volatility, already discussed, are also incorporated in the particular model.

While explaining the growth in market risk we made another, a more significant finding, in the context of European integration. Nowadays the situation (risk, volatility, etc.) in particular EMU stock markets is more affected by the general competitiveness of the sample of European economies. So the contemporary level of European integration already acknowledges the concept of "General Competitiveness of European Economy". In fact, the introduction of a single currency in EMU was another major step in this direction.

Table 4.
FGLS Estimates of joint model (eq. 22)
Dependent Variable: ϕ_i

Model	(3)	l (lag)
Constant Term	-0.661 (-0.495)	
General Competitiveness change	-4.27E-04*** (-2.683)	0
Individual Competitiveness change	-2.987E-03** (-2.262)	0
Change in domestic demand	0.021* (1.922)	3
Traded stock volume [c]	0.101 (2.074)**	1

Bond yields	-0.443*** (-5.871)	0
Foreign reserves [c]	0.331** (2.021)	1
Unemployement	0.118*** (3.554)	0
AR(1)	0.353*** (8.136)	
	Unweighted Statistics	
Adj. R-sq.	0.803	
S.E. of Regression	0.995	
	Significance of Group Effects Test	
F-stat	0.7288[a]	
F-crit. (1%)	2.6772	
	White General Test	
Chi-sq. stat	30.6952[b]	
Chi-sq. crit (1%)	30.5779	
	Included Observations	
Total panel obs.	480	
Obs. In cross sections	60	

Note:
a) $H_0 : b_{11} = ... = b_{n1}$ of common constant term is accepted.
b) H_0 of homoskedasticity is rejected.
c) Variables are expressed in logs.
t-stats. are given in the parentheses.
*** significant at 1%, ** significant at 5% confidence level. * significant at 10% confidence level.

5.2.2. Robustness checks

This section investigates the robustness of the empirical findings to a number of experiments with the estimated models (see appendix 3 tables 7-8). First, we tried the robustness of model one by one excluding the regressors. Signs and statistical significance are as expected, so that robustness with respect to EMU8 is not lacking. The other regressors are robust as well.

Next, a number of different lag structures were tried. We experiment with different lags for the regressors in the model (0, 3, 6, 9 and 12 month lags were tried one by one), to see how the EMU8 behaves. EMU8 is again robust. Coefficients and statistical significance for the other regressors in most cases also behave in an appropriate manner. However, in the case of change in domestic demand (TRADE), the coefficient keeps the positive sign for 3 and 6 month lag options, while the maximal significance is obtained for 3 month lag. Statistical significance of unemployment (UNEMPLOYMENT) lacks since 3 month lag and registers change in sign in the 6 month lag option. These cases can be interpreted as specific time limitations of the impact of these two factors and, in general, do not affect the robustness of the empirical model.

6. Conclusion

The stock markets of most EMU member states registered higher market risk after euro introduction. First of all, higher volatility of exchange rates affects the stock markets through consequent changes in the stock market value of firms. We show that exchange rates fluctuations affect the stock market risk by causing fluctuations in trade flows of the countries – our proxy for international competitiveness of the national economies.

Moreover, an even more interesting fact regarding this is that common currency strengthens the "net volatility" of changes in competitiveness for the entire sample of countries by synchronizing the changes of relative prices. Hence, the growth or reduction of Euro/USD exchange rates has a similar (positive or negative) effect on international competitiveness of all the economies of the Monetary Union (at least for the observed 8 member states).

The empirical study also shows that due to the deep economic integration of particular European economies at both governmental and corporate levels, the changes in "General competitiveness" are more significant in explaining the stock market risk in separate countries than the changes in competitiveness on national levels. This phenomenon indicates a new stage of European economic integration where a European corporations and brands are represented on the international market of goods and services.

Summarizing, the stock markets of most EMU member states registered higher market risk after euro introduction. Our analyses show that the Euro introduction had a triple effect on market risk, as it (1) resulted in higher volatility of exchange rates, (2) increased market risk on the stock markets because of higher synchronized fluctuations in general competitiveness, taking into account that (3) for the sample of countries it becomes more significant in explaining the dynamics of stock prices than the competitiveness changes at the national level.

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Appendix 1. Market risk dynamics

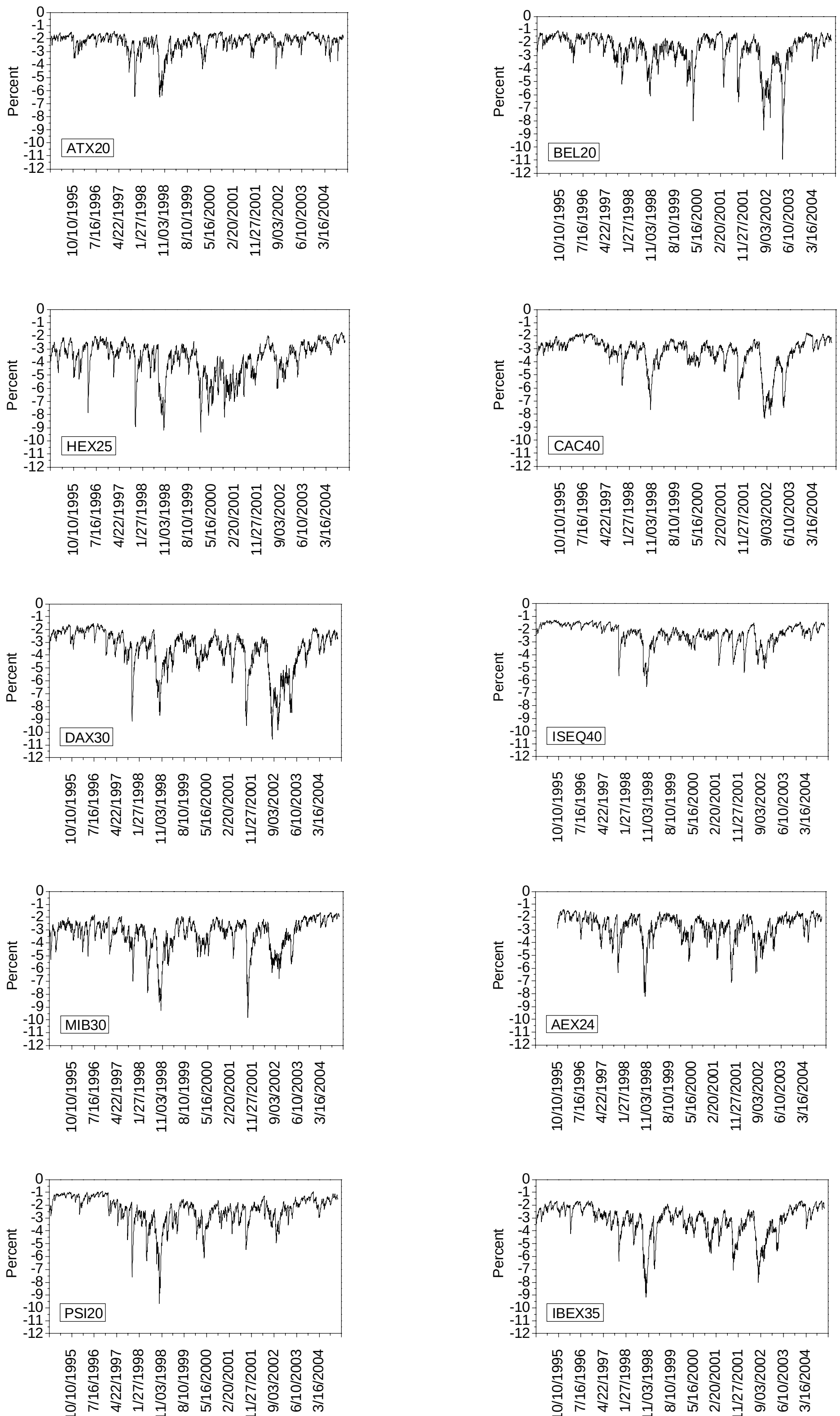

Figure 1. Value-at-Risk dynamics in EMU major stock markets: $VaR = -V\left(e^{\mu+\sigma\phi^{-1}(\alpha)} - 1\right)$, where V represents the initial value of some theoretical portfolio and $\phi(\cdot)$ is the cumulative distribution function of the standard normal probability distribution. GARCH (1.1) model is used for volatility forecasting.

Appendix 2. Data description

Table 5

Descriptive statistics for monthly data for the panel with 11 cross sections: 1995/01-2004/06

NET EXPORT

	Austria	Belgium	Finland	France	Germany	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
Mean	83.4	-390.1	104.0	628.2	2270.5	669.0	1036.3	-20.8	-890.0	40.1	-42.9
Median	55.1	-380.3	103.7	576.5	2241.3	454.2	1057.7	-2.5	-899.7	33.5	-54.5
Maximum	379.0	-22.9	270.4	1437.2	4269.5	2163.0	1759.4	16.1	-465.0	159.5	238.9
Minimum	-150.1	-693.1	-194.6	-32.4	753.0	-126.5	329.1	-226.5	-1213.9	-167.7	-325.8
Std. Dev.	103.7	142.6	70.8	336.6	861.4	629.1	287.3	54.4	179.3	45.5	111.9

GDP RATIO

	Austria	Belgium	Finland	France	Germany	Ireland	Italy	Luxembourg	Netherlands	Portugal	Spain
Mean	44.8	37.6	73.0	6.5	4.5	94.9	8.0	479.8	23.2	82.8	15.2
Median	44.9	38.0	74.8	6.7	4.5	96.3	8.0	473.3	23.8	82.9	15.4
Maximum	55.3	46.1	85.9	8.0	5.8	109.0	9.6	792.7	28.0	141.9	18.5
Minimum	30.4	25.5	55.0	4.6	2.9	78.0	6.2	412.4	17.0	66.6	12.2
Std. Dev.	8.0	6.5	9.7	1.1	0.9	7.7	1.1	62.4	3.1	13.0	1.7

REAL EXCHANGE RATE (EURO/USD)

Mean	111.7
Median	112.1
Maximum	141.3
Minimum	84.8
Std. Dev.	16.4

Note:

NET EXPORT	Net exports to USA (ml. USD) (ξ). Our own evaluations based on U.S. Census Bureau data
GDP RATIO	USA GDP/GDP $\left(\psi'/\psi\right)$ of the EMU member state ratio. Our own calculations based on Eurostat's quarterly data
REAL EXCHANGE RATE	Real exchange rates (ε) index (2000 average=100%). Source: Financial Statistics of the Federal Reserve Board and International Financial Statistics of the IMF.

Appendix 2. Data description (continued)

Table 6

Descriptive statistics for monthly data for the panel with 8 cross sections: 1998/10-2003/12

MARKET RISK

	Austria	Belgium	France	Germany	Italy	Netherlands	Portugal	Spain
Mean	2.248506	3.011755	3.820158	4.304242	3.745171	3.027104	3.00624	3.760452
Median	2.086519	2.617973	3.435668	3.812825	3.5486	2.837025	2.818665	3.482125
Maximum	4.973214	7.139443	7.469823	8.607236	7.730505	6.4128	7.326068	7.698977
Minimum	1.560709	1.24687	2.394064	2.330409	2.318071	1.889673	1.530123	2.163114
Std. Dev.	0.578273	1.312033	1.257402	1.603473	1.229161	0.932797	0.995916	1.177494

EMU8

Mean	-1.938133
Median	18.95272
Maximum	253.2226
Minimum	-380.4286
Std. Dev.	157.9432

MEMBER

	Austria	Belgium	France	Germany	Italy	Netherlands	Portugal	Spain
Mean	-0.054746	-0.134036	-0.386433	-0.925374	-0.284699	-0.070635	-0.024737	-0.057472
Median	0.535356	1.310721	3.778868	9.049102	2.784034	0.690725	0.241902	0.562013
Maximum	7.152762	17.51223	50.48853	120.9028	37.19679	9.228603	3.231992	7.50892
Minimum	-10.74594	-26.30946	-75.85134	-181.6381	-55.88252	-13.86457	-4.855576	-11.28101
Std. Dev.	4.461411	10.92295	31.49134	75.41103	23.20085	5.75618	2.015898	4.683558

TRADE

	Austria	Belgium	France	Germany	Italy	Netherlands	Portugal	Spain
Mean	2.247619	2.88254	4.265079	0.679365	2.261905	3.634921	4.260317	6.031746
Median	1.5	2.3	4	0.4	2.4	3.9	4.3	6
Maximum	13.6	9.8	10.3	6.2	5.4	10.8	16.6	10.5
Minimum	-3.6	-3.7	-0.7	-3.4	-1.1	-7.4	-7.9	1.9
Std. Dev.	3.633751	3.386717	2.150181	2.198985	1.253952	4.089529	4.63808	1.981106

LOG (TRADED)

	Austria	Belgium	France	Germany	Italy	Netherlands	Portugal	Spain
Mean	13.86582	15.04216	17.78988	17.82006	19.8233	15.20609	16.62162	18.40176
Median	13.8928	15.09747	17.93917	17.84267	19.8233	15.25649	16.86611	18.39836
Maximum	14.69503	15.88282	18.71098	18.61468	20.26482	16.03867	17.57519	19.18314
Minimum	13.12981	13.85015	16.58183	16.80993	19.12076	13.98976	14.98853	17.46229
Std. Dev.	0.350807	0.433255	0.68846	0.501226	0.23386	0.370557	0.679087	0.50653

BOND

	Austria	Belgium	France	Germany	Italy	Netherlands	Portugal	Spain
Mean	4.854603	4.894921	4.753016	4.649206	4.923492	4.766667	4.913651	4.862381
Median	5.06	5.08	4.93	4.78	5.13	4.92	5.09	5.05
Maximum	5.77	5.79	5.66	5.54	5.75	5.67	5.81	5.76
Minimum	3.74	3.74	3.69	3.62	3.82	3.72	3.77	3.69
Std. Dev.	0.578142	0.577514	0.538863	0.517096	0.556536	0.544311	0.581598	0.572867

Appendix 2. Data description (continued)

Table 6 (continued)

LOG (RESERVES)

	Austria	Belgium	France	Germany	Italy	Netherlands	Portugal	Spain
Mean	9.701383	9.526381	11.06698	11.40466	10.82213	9.852599	9.603671	10.55496
Median	9.768681	9.51392	11.05991	11.42412	10.86735	9.846864	9.634954	10.57457
Maximum	9.982128	9.907743	11.23022	11.51983	10.96809	10.19668	9.850219	11.06093
Minimum	9.21114	9.345133	10.89176	11.2474	10.59122	9.736133	9.224835	9.963123
Std. Dev.	0.211148	0.109276	0.097267	0.070559	0.10271	0.089106	0.157217	0.187303

UNEMPL

	Austria	Belgium	France	Germany	Italy	Netherlands	Portugal	Spain
Mean	3.933	7.573	9.360	8.454	9.844	3.048	4.776	11.617
Median	3.900	7.600	9.100	8.300	9.400	3.000	4.500	11.300
Maximum	5.100	9.600	11.400	10.300	11.800	4.400	6.500	15.000
Minimum	2.900	6.100	7.800	7.200	8.200	2.200	3.800	10.200
Std. Dev.	0.624	0.881	0.909	0.775	1.089	0.513	0.876	1.074

Note:

VaR	Stock market risk (%). VaR indicator is estimated for the indexes of particular EMU stock markets (ϕ). GARCH (1.1) model is used for the parameters estimation.
EMU8	Summed changes in net exports to USA for a sample of 8 EMU member states (Austria, Belgium, France, Germany, Italy, Netherlands, Portugal and Spain) caused by the changes of real exchange rates (ml. USD). Source: Our own evaluations based on U.S. Census Bureau data $\left(\sum_{i=1}^{8}\Delta\varepsilon_{i(t-5-l)}b_{i1}\right)$.
MEMBER	Changes in net exports to USA of particular EMU member state caused by the changes of real exchange rates (ml. USD). Source: Our own evaluations based on U.S. Census Bureau data $\left(\Delta\varepsilon_{i(t-5-l)}b_{i1}\right)$.
TRADE	Monthly growth rates of retail trade $(\Delta\lambda)$ compared to the same period of the previous year (%). Source: Eurostat.
TRADED	Traded volume of stocks. Source Reuters. (ν).
BOND	Long-term government bond yields (γ) (monthly average, not seasonally adjusted). Source: Eurostat.
RESERVES	Foreign official reserves, including gold in million euros (end of period). Source: Eurostat.
UNEMPL	Harmonised unemployment rates (ρ).Unemployment according to ILO definition (%). Source: Eurostat.

Appendix 3. Robustness checks

Table 7

Excluding regressors

Number of regressors excluded from equation	(0)	(1)	(2)	(3)	(4)	(5)
EMU8	-0.0003 (-2.6466)	-0.0003 (-2.5418)	-0.0002 (-2.2362)	-0.0002 (-1.6357)	-0.0002 (-1.7715)	-0.0002 (-1.7811)
TRADE	0.0162 (2.7311)	0.0157 (2.5552)	0.0149 (2.4561)	0.0088 (1.3984)	0.0118 (1.8539)	
LOG(TRADED)	0.1278 (1.8941)	0.2029 (3.2333)	0.2799 (5.1262)	0.2999 (5.1754)		
BOND	-0.4018 (-4.6234)	-0.4389 (-4.8851)	-0.3765 (-4.3043)			
LOG(RESERVES)	0.3748 (1.6841)	0.5289 (2.3873)				
UNEMPLOYMENT	0.1429 (2.7170)					
Adj. R2	0.6044	0.6056	0.6077	0.5989	0.5866	0.5876

Note: t-stats. are given in parentheses.

Appendix 3. Robustness checks (continued)

Table 8

Changing the lags for the regressors

Lags	(0)	(3)	(6)	(9)	(12)
EMU8*	-0.0003 (-2.3723)	-0.0002 (-1.5580)	-0.0003 (-2.0738)	-0.0002 (-1.4863)	-0.0002 (-1.5865)
TRADE	-0.0082 (-1.3247)	0.0125 (1.8725)	0.0092 (1.3306)	-0.0005 (-0.0620)	0.0145 (2.0475)
LOG(TRADED)	0.1470 (2.0918)	0.1514 (2.1711)	0.2088 (2.9526)	0.1481 (2.1064)	0.1745 (2.4948)
BOND	-0.3335 (-3.4699)	-0.3274 (-3.3469)	-0.1923 (-1.8420)	-0.1796 (-1.6741)	-0.0392 (-0.3626)
LOG(RESERVES)	0.3628 (1.5528)	0.4799 (1.8887)	0.6282 (2.3427)	0.7376 (2.8009)	0.8915 (3.2278)
UNEMPLOYMENT	0.1372 (2.5370)	0.0649 (1.1745)	-0.0013 (-0.0231)	0.0365 (0.6747)	0.0173 (0.3135)
Adj. R2	0.6056	0.6049	0.6047	0.6034	0.6033

Note: * lag is kept invariant as it appears in the original model.
t-stats. are given in the parentheses.

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