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ENHANCING FORECAST ACCURACY BY USING LONG ESTIMATION PERIODS

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ABSTRACT

A tradeoff between forecast accuracy and the length of an estimation period always exists in forecasting. Longer estimation periods are argued to be less efficient, however, using the forecast encompassing and accuracy test, this study discusses the importance of considering the overall usefulness of information in the in-sample period. The empirical results demonstrate that forecasts using the correct model have reduced measurement loss and the mean of forecast errors decrease with an increase in in-sample period. Moreover, for the forecast accuracy and encompassing tests, reducing the use of observations in making estimates leads to the wrong model being easily accepted. Additionally, these analytical results are also consistent with the application in hedge performance, that is, the hedge effectiveness is optimized when the estimation period is longest, particularly under the recursive scheme.

INTRODUCTION

Improving out-of-sample forecast accuracy is a key concern in numerous areas of economics and finance. Increasingly sophisticated economic models are being developed to fit real time series, but surprisingly a good in-sample fit does not necessarily translate into good out-of-sample performance. This surprising occurrence is due to model instability. Sample period length is one of the foundations of accurate in-sample estimates and out-of-sample forecasts. Forecasting agents often consider it necessary to estimate model parameters using only a partial window of the available observations, to avoid heterogeneity within the results. Some researches argue that a longer sample period achieves an increase in forecast bias and thus reduces efficiency. Shaffer (2003) notes that too large a sample may include older observations that may reflect different biases to those existing more recently. Harris and Shen (2003) proposed that a longer estimation period is associated with less efficient estimations when compared with short estimation periods. However, Clark and McCracken (2004) notes that reducing the sample, to lower the heterogeneity, also increases the variance of the parameter estimation, and increases forecast errors and mean square forecast errors (MSE). Therefore, the decision whether or not to use all available data when constructing a forecast is problematic and often results in a tradeoff. However, little attention has been given to addressing the problem of determining appropriate estimation periods.

LITERATURE REVIEW

The forecast approaches are also related to this point and two other common approaches. Both the rolling window approach and the recursive scheme, are considered here for one-step-ahead forecasting and provide evidence for the argument above. In the rolling scheme the forecasting model is estimated using a moving window of recent observations as the forecast progresses through time. “Rolling windows” is a commonly used concept in financial literature. Giacomini and White (2005) argue that a rolling window approach with limited memory estimators is appropriate in heterogeneous data environments. Many studies adopt a rolling window approach to forecast. Clements and Hendry (1988, 1999) propose that economic time series are often heterogeneous. However, this study argues that even in the rolling window scheme, the problem of estimate bias still exists, when the window size is too short to contain enough useful information. Otherwise, the recursive scheme indicates that estimating with more data as forecasting moves forward in time, and this expanding window approach is generally applied in the macroeconomic literature uses all available data (Stock and Watson, 2003). In a stationary environment, the recursive scheme is necessary since the limited memory estimators are inefficient (Giacomini and White, 2005). Therefore, as noted by Pesaran and Timmerman (2004), a forecasting approach that is based on the breakpoint risk must be inefficient as it does not fully use all available information. Regardless of the environment or forecast approach used, sufficient useful information should be contained in the estimation periods to reduce the bias parameter estimate and improve the out-of-sample forecast performance.

This study considers two sets of out-of-sample forecasts and analyses the relationship between spot and nearby futures of the West Texas Intermediate (WTI) crude oil price. The study encompasses both those with and without structural change. The spot price increased 51.93% from January to August, 2005[1]. Forecasting the relationship is not only relevant to oil market traders, but also to global economic activity and government policy. Bai and Perron (1998, 2003)[2] have already demonstrated that structural change exists in the time series data, and eliminating this structural change will bias the forecast results. Moreover, such biases can accumulate and produce larger mean square forecast errors (see, Clark and McCracken, 2004; Jardet, 2004; Inoue and Kilian, 2002; Chauvet and Potter, 2002; Krolzig, 2001; Koop and Potter, 2000; Clements and Hendry, 1999). Consequently, this study designates the model with a structural break as the correct model, and the model without such a structural break as the incorrect model. Intuitively one may hold the belief that the out-of-sample forecast error under the correct model will be lower than that under the incorrect model, but interestingly this study finds that the results are adverse given a shorter estimation period. Longer estimation periods including full information not only produce consistent parameter estimations, but also improve the forecast performance and hedge effectiveness.

Moreover, this investigation adopted a new test to assess the reliability of the out-of-sample forecast abilities regarding the usefulness of the information contained for forecasting, termed the forecast encompassing test[3] (Clark and McCracken, 2001, 2004). As previously stated, mean square forecast errors are the most widely used criterion for testing forecasting abilities (for example, Stock and Watson, 2003 and Hamilton, 2001). However, recently forecast encompassing tests have provided a possible means of complementing the MSE criterion (Rapach and Weber, 2004). The tests effectively reveal whether one variable can be used to predict another. Clark and McCracken (2001) propose that the encompassing tests are the most effective for post-sample testing. For intuitive forecasting results, the model with a structural break should have a smaller mean square forecasting error and should encompass another model based on the better use of forecasting information. However, the empirical results differ when the estimation periods are insufficiently long, and the situation is modified with increasing period length.

The remainder of this paper is organized as follows: Section 2 describes the empirical methodology, including the nested model-setting and the tests of forecast accuracy and encompassing. Section 3 presents data and the empirical model for considering structural change. The application in hedge effectiveness is also drawn upon in this section. Section 4 describes the empirical results. Section 5 draws conclusions on the evidence provided in the paper.

METHODOLOGY

This study uses a simple nested linear model, with or without structural change, to demonstrate that one-step-ahead forecasting[4] with longer estimation periods performs better when more useful information is contained. Both rolling window and recursive schemes are used to examine the forecast accuracy and the encompassing nature of different estimation periods. Additionally, hedge performance is also considered in this investigation.

Nested Model-Setting and Forecast Scheme

Following Clark and McCracken (2001, 2004), sample of observations $\{y_t, x'_{2,t}\}_{t=1}^{T+1}$ contains a random variable y_t to be forecast and a $(k_1 + k_2 = k \times 1)$ vector of predictors $x_{2,t} = (x'_{1,t}, x'_{22,t})'$. $x_{1,t}$ denotes the regressors in the restricted model (model 1) and $x_{2,t}$ represents the regressors in the unrestricted model (model 2) with k_2 variables. The in-sample observations span 1 to R, and the out-of-sample observations span $R+1$ to $R+P$, where P denotes the number of one-step ahead forecasts. The total number of observations in the sample is $T+1=R+P$. Moreover, forecasts of y_{t+1}, $t = R, \cdots T$ are generated using two linear models with the form $x'_{i,t+1}\beta_i^*$, $i=1,2$, each of which is estimated. Under the null hypothesis, model 2 nests model 1, and thus model 2 includes k_2 excess parameters. Under the alternative hypothesis, the restriction k_2 is not true and model 2 is correct.

The forecasting schemes are permitted to be both rolling and recursive one-step ahead predictions. Under the rolling scheme, forecasting models are estimated using a moving window of the most recent R observations as the forecast horizon moves further forward. Under the recursive scheme, forecasting models are estimated using more data as the forecast horizon moves further forward, and the maximum number of observations used for parameter estimation is $T = R + P - 1$.

Forecast Accuracy and Forecast Encompassing Tests

This study denotes the one-step ahead forecast errors as $\hat{u}_{1,t+1} = y_{t+1} - x'_{1,t}\hat{\beta}_{1,t}$ and $\hat{u}_{2,t+1} = y_{t+1} - x'_{2,t}\hat{\beta}_{2,t}$ for model 1 and model 2, respectively. Clark and McCracken (2001) treat the tests for equal MSE (MSE-T and MSE-F) and forecast encompassing (ENC-T and ENC-F) as one-sided tests. The asymptotic distributions of equal MSE tests are derived by McCracken (2004), an F-type test proposed by McCracken (2001) (MSE-F), and a T-test developed by Diebold and Mariano (1995) and West (1996) (MSE-T). The test of forecast encompassing is derived by Clark and McCracken (2001). The statistics on tests regarding MSE equality (also called the 'forecast accuracy test') are described simply as follows:

$$MSE - T = P^{1/2} \frac{P^{-1}\sum_{t=R}^{T}(\hat{u}_{1,t+1}^2 - \hat{u}_{2,t+1}^2)}{\sqrt{P^{-1}\sum_{t=R}^{T}(\hat{u}_{1,t+1}^2 - \hat{u}_{2,t+1}^2)^2}} \tag{1}$$

$$MSE - F = P \times \frac{P^{-1}\sum_{t=R}^{T}(\hat{u}_{1,t+1}^2 - \hat{u}_{2,t+1}^2)}{P^{-1}\sum_{t=R}^{T}\hat{u}_{2,t+1}^2} \tag{2}$$

The null hypothesis is that the MSE of model 1 is less than or equal to that of model 2, while the alternative hypothesis is that the MSE of model 1 exceeds that of model 2. Furthermore, in the test for forecast encompassing, the null hypothesis is that the forecast produced with model 1 encompasses model 2, the covariance in the numerator of the encompassing tests statistics will be less than or equal to 0. The alternative hypothesis is that model 2 includes more information and the covariance should be positive. The statistics are reported as follows:

$$ENC-T = P^{1/2}\,\frac{P^{-1}\sum_{t=R}^{T}(\hat{u}_{1,t+1}^{2}-\hat{u}_{1.t+1}\hat{u}_{2,t+1})}{\sqrt{P^{-1}\sum_{t=R}^{T}(\hat{u}_{1,t+1}^{2}-\hat{u}_{1.t+1}\hat{u}_{2,t+1})^{2}}} \tag{3}$$

$$ENC-F = P\times\frac{P^{-1}\sum_{t=R}^{T}(\hat{u}_{1,t+1}^{2}-\hat{u}_{1,t+1}\hat{u}_{2,t+1})}{P^{-1}\sum_{t=R}^{T}\hat{u}_{2,t+1}^{2}}\quad . \tag{4}$$

Data and the Empirical Model

This study developed a nested linear model for analyzing the relationship between spot and nearby futures of West Texas Intermediate (WTI) crude oil price. The available data, obtained from the U.S. Department of Energy, is for the period June 23, 1988 to June 28, 2005 and includes a total of 4,178 observations. The restricted (model 1) and unrestricted models (model 2) are constructed as Eqns. (5) and (6):

$$S_t = \alpha_0 + \beta_0 F_t + u_{1t} \tag{5}$$

$$S_t = (\alpha_0 + \alpha_1 D_t) + (\beta_0 + \beta_1 D_t)\times F_t + u_{2t} \tag{6}$$

Where S_t and F_t are the continuously compounded returns of the WTI crude oil spot and nearby futures prices. The variable D_t equals one when the date is after the break and otherwise equals zero. The variables u_{1t} and u_{2t} represent the error terms for models 1 and 2, respectively. According to Bai and Perron (2003), one structural break was obtained in March 21, 1997, indicating that the coefficient shifted over the break date.

The out-of-sample periods run from the break date to the end of the data (March 21, 1997 – June 28, 2005), and the in-sample periods differ depending on the observation period length. Taking 500 days as an example, the first in-sample period begins at 500 days before the structural break date is taken to be March 9, 1995 to March 20, 1997. The second round forecast starts at 499 days before the structural date, that is March 10, 1995 under the rolling window scheme, while the starting date of the estimation period is fixed at March 9, 1995 under the recursive scheme. This study considers three estimation periods; one-year (250 days), two-years (500 days) and five-years (1250 days) under the rolling scheme. For the recursive scheme the setting is different. Due to the increase in the number of observations as the time of the forecast moves forward, the start points of the estimation periods are only set to either 250 or 500 days before the break date and the number of observations considered in forecasting increases as the time moves forward.

Application in Hedge Effectiveness

Additionally this investigation assessed the hedge effectiveness to examine the improvement in effectiveness with increasing estimation periods. While facing an oil market characterized by high volatility, eliminating or lowering risk used to be a key objective of majority traders with futures contracts being the most widely used method of achieving this. Hedge effectiveness can be defined as the proportion of the variance that is eliminated by the hedge, with hedge effectiveness and performance increasing with the size of the reduction. The unhedged variance is expressed as

$$Var(U) = \sigma_u^2 = Var(S_t)\ , \tag{7}$$

where S_t denotes the continuously compounded spot price returns. Alternatively, the hedged variance is calculated from the hedged return, which can be written as follows:

$$X_t = S_t - hr_t F_t\ , \tag{8}$$

where F_t represents the continuously compounded futures price returns, the coefficient hr_t is the hedge ratio which is known as the coefficient β in equation (5) and (6), and X_t denotes the return of the hedge investment. Thus, the hedged variance of the hedge equation is expressed as

$$Var(H) = \sigma_h^2 = Var(X_t) \tag{9}$$

The hedge effectiveness (HE) can be assessed using Eqn. (10). Hedge performance improves with increasing HE.

$$HE = \frac{Var(U) - Var(H)}{Var(U)} = \frac{\sigma_u^2 - \sigma_h^2}{\sigma_u^2} \tag{10}$$

EMPIRICAL RESULTS

Under the Rolling Scheme

Table 1 lists the out-of-sample forecasts. Notably, the MSE decreases with an increasing in-sample period under the correct model in this study. However, when the observation period is short, the forecasting errors under the correct model are not lower than under the incorrect one. For example, during the 250 day estimation period, the MSE is found to be 1.5926 in model 1, which exceeds the 1.5943 in model 2. The test statistics for MSE equality are significantly negative under the 1% level, indicating that the forecast error in model 2 is significantly higher than in model 1, the relationship between spot and nearby futures of WTI crude oil price, that is, the real model performs worse in one-step-ahead forecasting. Nevertheless, the situation reverses when the estimate period is increased to either 500 or 1250 days, and the MSE in unrestricted. Model 2 becomes significantly smaller than in the restricted model 1. This indicates that more useful information during the estimate period will reduce the variance of parameter estimation, and increase the accuracy of the out-of-sample forecast. Obviously, reducing the sample increases the variance of the parameter estimation and may even obtain reverse results.

Stronger evidence comes from forecast encompassing tests. In both the 500 and 1250 day estimation periods, the null hypotheses of model 1 encompassing model 2 are significantly rejected at the 1% levels. This means that the model with structural change contains more information than the model without structural change. However, the results for the 250 day estimation period are not as clear as for longer periods, and the conclusions reached are inconsistent between the 250 day estimation period and the longer estimation periods. The statistics in ENC-T tests is 2.642, and is significant at the 1% level, and the statistics in ENC-F is 3.767, indicating that there is insufficient evidence to reject the assumption that the dummy variable is useful. According to Clark and McCracken (2001, 2004), ENC-F is the most powerful measure, followed by ENC-T and MSE-F, while the least powerful measure is MSE-T. Therefore, without the powerful support provided by ENC-F, it cannot be concluded that model 2 does have more useful information than model 1 in 250 day estimation periods. Obviously, this is inconsistent with the fact that the structural break already exists in this instance.

Table 1: Empirical Results under the Rolling Scheme

Model 1: $S_t = \alpha_0 + \beta_0 F_t + u_{1,t}$				Model 2: $S_t = \alpha_0 + \alpha_1 D_t + (\beta_0 + \beta_1 D_t)F_t + u_{2,t}$			
	250 days	500 days	1250 days		250 days	500 days	1250 days
MAE	0.6340	0.6403	0.6392	MAE	0.6335	0.6326	0.6283
RMSE	1.2620	1.2656	1.2635	RMSE	1.2626	1.2621	1.2606
MSE	1.5926	1.6017	1.5966	MSE	1.5943	1.5930	1.5893
Tests for equal MSE (Forecast accuracy test)							
	250 days	500 days	1250 days				
MSE-F	-2.154**	11.111**	9.344**				
MSE-T	-0.753*	1.133**	0.688*				
Forecast encompassing test							
	250 days	500 days	1250 days				
ENC-F	3.767	16.879**	17.827**				
ENC-T	2.642**	3.453**	2.655**				

*Notes: Model 1 and model 2 are restricted and unrestricted models, respectively. MAE is the mean of the absolute forecast error, MSE is the mean square error, and RMSE is the square root of MSE. The forecast accuracy test and forecast encompassing test statistics and the relative critical values are suggested by Clark and McCracken (2001, 2004). **, * represent significances under the 1% and 5% levels.*

Under the Recursive Scheme

Table 2 lists the outcomes achieved using the recursive scheme. The MSE of model 2 remains almost unchanged with an increasing estimation period. This pattern is different from the rolling scheme, the recursive scheme is characterized by increasing the information added as forecasting moves forward in time. Therefore, when the forecast horizon becomes far enough away, the forecast errors become almost identical owing to almost identical sized sets of information being used for the estimation. As shown in this study, the forecast horizon is from March 21, 1997 to June 28, 2005. The horizon contains a total of 2,023 forecast days, and the MSEs are almost fixed at 1.5885 in any in-sample period. Besides, in the correct model-setting, forecasting under the recursive scheme obtains smaller MSE than under the rolling scheme resulting from the overall information involved.

Regarding the forecast accuracy test, MSE in model 2 are significantly lower than in model 1 at the 1% level for the 500 estimation period. However, for the 250 day period, the results of the MSE-F and MSE-T tests are inconsistent. The value of MSE-F is 17.474 and does not provide enough evidence to reject the null hypothesis that MSE is identical in models 1 and 2, but the MSE-T test is significantly below the 1% level. Based on the suggestion by Clark and McCracken (2001, 2004), the forecast accuracy tests are less useful than the forecast encompassing tests, thus, this study more closely examines the ENC tests. As for the forecast encompassing tests, the results are consistent among estimation periods. All the statistics are significant under the 1% level, indicating that model 2 really contains more useful information than model 1. The dummy variable of structural change is necessary, and eliminating the feature of structural change would increase the forecasting error.

Hedge Effectiveness

The empirical results are listed in Table 3. Under either rolling or recursive schemes, the hedge effectiveness is higher in model 2 except for the short estimation period, 125 days. The hedge performance is said to be improved in the correct model, but the result is biased during the short estimation period. Furthermore, under the rolling scheme with the correct model, the values of hedge effectiveness are 0.77111 and 0.77144 for the 500 day and 1250 day estimation periods respectively. The hedge performance improves with increasing estimation period. Compared with the recursive scheme, the hedge effectiveness is around 0.7716, and all values are higher than for the rolling scheme in model 2. In summary, under the condition of the model with structural change, the hedge effectiveness is optimized when the estimation period is longest, particularly under the recursive scheme. These results are consistent with the above arguments; that is, the hedge performance is better under the recursive scheme owing to the consideration of all information.

Table 2: Empirical Results under the Recursive Scheme

Model 1: $S_t = \alpha_0 + \beta_0 F_t + u_{1,t}$			Model 2: $S_t = \alpha_0 + \alpha_1 D_t + (\beta_0 + \beta_1 D_t)F_t + u_{2,t}$		
	250 days	500 days		250 days	500 days
MAE	0.6454	0.6494	MAE	0.6271	0.6271
RMSE	1.2657	1.2671	RMSE	1.2604	1.2603
MSE	1.6022	1.6054	MSE	1.5885	1.5885
Tests for equal MSE (Forecast accuracy test)					
	250 days	500 days			
MSE-F	17.474	21.582**			
MSE-T	1.199**	1.317**			
Forecast encompassing test					
	250 days	500 days			
ENC-F	26.624**	31.714**			
ENC-T	3.637**	3.842**			

*Notes: Model 1 and model 2 are restricted and unrestricted models, respectively. MAE is the mean of the absolute forecast error, MSE is the mean square error, and RMSE is the square root of MSE. The forecast accuracy test and forecast encompassing test statistics and the relative critical values are suggested by Clark and McCracken (2001, 2004). **, * represent significances under the 5% and 1% levels.*

Table 3: Hedge Effectiveness

	Rolling scheme		Recursive scheme	
Estimation period	Model 1	Model 2	Model 1	Model 2
250 days	0.77132	0.77130	0.76926	0.77159
500 days	0.76954	0.77111	0.76874	0.77158
1250 days	0.76999	0.77144		

CONCLUSION

The length of estimation period is always a problem in forecasting. An excessively long in-sample period is charged with reducing forecast efficiency, while an excessively short sample period will increase the variance of the parameter estimates and bias the out-of-sample forecasts. Accordingly, this study uses a simple nested linear model to demonstrate that one-step-ahead forecasting with longer estimation periods performs better when more information is contained. Both rolling window and recursive schemes are used to examine the forecast accuracy and encompassing for different estimation periods. The empirical results show that forecasts under the correct model reduces measurement loss, and the mean square forecast errors decrease with increasing in-sample period. The inclusion of more information in the estimate period lowers the variance of parameter estimation, and increases the accuracy of the out-of-sample forecast. For the forecast accuracy and encompassing tests, the use of fewer observations in making an estimate could easily lead to wrong decisions and the acceptance of the wrong model. Finally, these results are also consistent with hedge effectiveness, namely that the effectiveness is better under the recursive scheme in terms of considering all useful information.

END NOTES

[1] The oil price was $43.96 per barrel on 4 January, 2005 and was $66.79 per barrel on August 15, 2005.

[2] Gabriel et al. (2003) note that testing for structural change is a means of testing the model specifications.

[3] The preferred forecasts, namely those with better performance, depend on the competing forecasts lacking information. Chong and Hendry (1986) and Clements and Hendry (1993) termed this situation the preferred forecasts encompassing the competing forecasts. Clark and McCracken (2001, 2004) developed and formulated the tests.

[4] Harvey et al. (1998) suggested that it is reasonable to assume that the forecast errors of one-step-ahead forecasts are not autocorrelated, so that the regression-base test is very straightforward to implement.

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THE HESTON STOCHASTIC VOLATILITY MODEL FOR SINGLE ASSETS AND FOR ASSET PORTFOLIOS: PARAMETER ESTIMATION AND AN APPLICATION TO THE ITALIAN FINANCIAL MARKET

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ABSTRACT

We investigate the performance of the Heston stochastic volatility model in describing the probability distribution of returns both in the case of single assets and in the case of asset portfolios. The R. parameters of the Heston model are estimated from observed market prices using a simple calibration method based on an integral representation of the exact probability density function of returns derived by Dragulescu and Yakovenko (2002). In the case of multiple correlated assets, the correlation parameters are obtained using a heuristic procedure based on a matrix completion algorithm. We present numerical experiments where several stocks traded on the Italian financial market are considered. We show that, both in the case of single assets and in the case of multiple correlated assets, the Heston model provides an excellent agreement with historical time series data and fits the empirical probability distribution of returns far better than the lognormal model.

INTRODUCTION

In this paper we assess the performances of the Heston model (HM) in describing the probability distribution of stock returns on the Italian financial market. At the same time we propose a simple method to calibrate the HM that gives an excellent agreement with historical time-series data both in the case of single and multiple correlated assets.

The paper is organized as follows: in Section 2 we provide a review of related literature. In Section 3 we briefly recall the basic facts about the LM. In Section 4 we give a description of the HM in the case of single and multiple correlated assets. In Section 5 we describe the calibration method used to estimate the parameters of the HM. Finally, in Section 6 we present and discuss the numerical results obtained applying the calibration algorithm developed in Section 5.

LITERATURE REVIEW

The dynamics of stock market prices is often described by the lognormal model (LM). Based on the assumption of constant drift and volatility, the LM gives a normal probability distribution of asset returns and therefore it is a very simple and tractable model. This is the reason why the LM, originally introduced by Bachelier (1900) and refined by Osborne (1959), is still nowadays very popular among financial researchers and practitioners.

Nevertheless many empirical studies on financial markets show that the probability distribution of stock returns is far from being normal. In particular, empirical observations of option prices reveal that the volatility of the underlying stocks varies as a function of the strike prices (the so-called smile effect, see Wilmott, 1998). Moreover the probability distribution of realized returns is often leptokurtic, i.e. it has fatter tails and higher peaks than the lognormal probability distribution (Bouchaud & Potters, 2001, Fama, 1965).

This empirical evidence motivated several authors to reject the assumption of constant volatility and to introduce the so-called stochastic volatility models; that is, models where the asset price volatility is described as a stochastic process. Among the stochastic volatility models that can be found in the literature, see for instance Heston, 1993, Hull & White, 1987, Melino & Turnbull, 1990, Scott, 1987, Stein & Stein, 1991, the Heston model (Heston, 1993) has received considerable attention since it gives an adequate description of stock market dynamics (Dragulescu & Yakovenko, 2002, Prange, Silva & Yakovenko, 2004) and yields tractable closed-form solutions (Dragulescu & Yakovenko, 2002, Heston, 1993).

In the financial literature the calibration of the HM is usually performed using two different approaches: in Aboura, 2004, Forbes, Martin & Wright 2002, Heston, 1993, Pan, 2002, the HM is calibrated consistently with observed option prices, while in Daniel, Joseph & Bree, 2005, Dragulescu & Yakovenko, 2002, Prange, Silva & Yakovenko 2004, Silva & Yakovenko, 2001 the parameters of the HM are estimated by fitting the probability distribution of realized asset prices.

In this paper we follow the latter approach since several stocks traded on the Italian financial market do not have options written on them. In particular, in the case of single assets, the calibration method proposed in this manuscript is similar to the one developed by Dragulescu and Yakovenko (2002). In Dragulescu & Yakovenko, 2002, an integral representation of the probability density function of returns of the HM is derived (Formula (23) p. 446). Note that this formula is an exact formula that gives the probability distribution of returns conditioned to the value taken by the initial variance. In addition Dragulescu and Yakovenko (2002) obtain also another expression for the probability density function of returns of the HM (Formula (28) p. 446), where the initial variance does not appear. This second formula is an approximate formula, since it is based on the assumption that the probability distribution of the initial variance is equal to the steady state probability distribution of the variance process.

The approximate formula of Dragulescu & Yakovenko (2002) is used in Daniel, Joseph & Bree, 2005 to calibrate the HM against the Dow Jones Industrial Average, S&P 500, FTSE 100 indexes, in Dragulescu & Yakovenko, 2002 to estimate the parameters of the HM against Dow Jones index time-series data, in Prange, Silva & Yakovenko, 2004 to estimate the parameters of the HM for several stocks belonging to the Dow Jones index, and in Silva & Yakovenko, 2001 to calibrate the HM for the S&P 500, NASDAQ and Dow Jones indexes. Note that in these works the value taken by the initial variance of the asset returns is not estimated, since it is not contained in the approximate formula of Dragulescu & Yakovenko, 2002.

The calibration method developed in this paper is based on the exact closed-form expression of the probability density function of the HM derived in Dragulescu & Yakovenko, 2002, Formula (23) p. 446. In particular we treat the initial variance of asset returns as an additional parameter of the model, so that we do not have to assume that the initial variance has stationary probability distribution, as done in Daniel, Joseph & Bree, 2005, Dragulescu & Yakovenko, 2002, Prange, Silva & Yakovenko, 2004, Silva & Yakovenko, 2001. Moreover, contrary to the calibration methods proposed in Daniel, Joseph & Bree, 2005, Dragulescu & Yakovenko, 2002, Prange, Silva & Yakovenko, 2004, Silva & Yakovenko, 2001, our approach allows to estimate the value taken by the initial variance.

First of all the method to calibrate the HM is developed in the case of single assets and then it is extended to the case of multiple correlated assets. In this latter case the correlation parameters of the HM are estimated using a heuristic technique based on a suitable matrix completion algorithm.
We estimate the parameters of the HM for several stocks belonging to the Italian Stock Exchange using historical data on a daily basis from June 2002 to June 2006. The results obtained show that using the calibration algorithm proposed in this paper the HM provides an excellent agreement with empirical data

both in the case of single and multiple correlated assets. In particular the HM captures the kurtosis effect exhibited by the empirical probability distribution of returns.

We point out that the contribution of this paper is twofold. First, we show that the HM describes the probability distribution of returns far better than the LM for stocks belonging to the Italian financial market. Second, we propose a method to calibrate the HM that is very easy to implement and performs very well both in the case of single assets and in the case of asset portfolios. From the practical standpoint we believe that these results can be very interesting for a stock trader. In fact it is crucial for a financial investor to use a mathematical model of stock prices which is simple to calibrate and provides a good agreement with realized returns.

THE MODEL

The Lognormal Model

Let $S(t)$ denote the price of an asset at time t and let t_0 denote the current time. According to the lognormal model (LM), $S(t)$ is described as a stochastic process satisfying the stochastic differential equation:

$$\frac{dS(t)}{S(t)} = \mu\, dt + \sigma dW(t), \quad t \geq t_0, \tag{1}$$

with initial condition:

$$S(t_0) = S_0. \tag{2}$$

In (1), μ and σ are constant parameters, called drift and volatility, respectively, and $W(t)$ is a standard Wiener process.
Let us define the asset return over the time interval $[t_0, t]$:

$$X(t) = \log\left(\frac{S(t)}{S_0}\right), \quad t \geq t_0. \tag{3}$$

Using Ito's lemma, equation (1) and initial condition (2) can be rewritten as follows:

$$dX(t) = \left(\mu - \frac{\sigma^2}{2}\right)dt + \sigma dW(t), \quad t \geq t_0, \tag{4}$$

$$X(t_0) = 0. \tag{5}$$

The parameters μ and σ can be estimated from realized asset prices in a very simple way. In fact let us consider a set of equally spaced time values t_0, t_1, …,t_n and let us define $\Delta t = t_k - t_{k-1}$, $k=1,2,\dots,n$. Equation(4) implies that asset returns are distributed as follows:

$$\log\left(\frac{S(t_k)}{S(t_{k-1})}\right) \approx Normal\left(\left(\mu - \frac{\sigma^2}{2}\right)\Delta t, \sigma\sqrt{\Delta t}\right), \; k = 1,2,\dots,n, \tag{6}$$

where $Normal(a,d)$ denotes the normal probability distribution with mean a and standard deviation d.

Let S_k denote the asset price observed at time t_k, k=1,2,…,n, let us consider the realized returns:

$$x_k = \log\left(\frac{S_k}{S_{k-1}}\right),\ k = 1,2,\ldots,n, \tag{7}$$

and let *mean* and *var* respectively denote the sample mean and the sample variance of the realized returns:

$$mean = \frac{1}{n}\sum_{k=1}^{n} x_k, \quad var = \frac{1}{n-1}\sum_{k=1}^{n}(x_k - mean)^2. \tag{8}$$

The probability distribution of returns (6) yields the following statistical estimators for the parameters σ and μ:

$$\sigma = \sqrt{\frac{var}{\Delta t}}, \quad \mu = \frac{mean}{\Delta t} + \frac{\sigma^2}{2}. \tag{9}$$

Now let us consider the LM in the case of multiple correlated assets. Let m denote the number of assets considered and let $S_k(t)$ denote the price of the i-th asset at time t, i=1,2,…,m. Let us define the asset returns:

$$X_i(t) = \log\left(\frac{S_i(t)}{S_i(t_0)}\right),\ t \ge t_0,\ i = 1,2,\ldots,m, \tag{10}$$

equation (4) and initial condition (5) are generalized to the case of m correlated assets as follows:

$$dX_i(t) = \left(\mu_i - \frac{\sigma_i^2}{2}\right)dt + \sigma_i dW_i(t),\ t \ge t_0,\ i = 1,2,\ldots,m, \tag{11}$$

$$X_i(t_0) = 0,\ i = 1,2,\ldots,m, \tag{12}$$

where μ_i and σ_i represent the drift and the volatility of the i-th asset, respectively, and $W_i(t)$ is a standard Wiener process, i=1,2,…,m. The correlation coefficient between $W_i(t)$ and $W_j(t)$, which we denote with $\rho_{i,j}$, is assumed to be constant, i=1,2,…,m, j=1,2,…,m.

According to equation (11) the vector of stochastic variables [$X_1(t)$, $X_2(t)$, …, $X_m(t)$] has a multivariate normal distribution. Therefore in analogy with the case of single assets the parameters μ_i, σ_i, $\rho_{i,j}$, i=1,2,…,m, j=1,2,…,m, can be estimated from observed asset prices as follows.
Let $x_{i,k}$ denote the realized return of the i-th asset over the time interval [t_{k-1}, t_k], i=1,2,…,m, k=1,2,…,n. Let us define:

$$mean_i = \frac{1}{n}\sum_{k=1}^{n} x_{i,k}, \quad var_i = \frac{1}{n-1}\sum_{k=1}^{n}(x_{i,k} - mean_i)^2,\ \ i = 1,2,\ldots,m, \tag{13}$$

$$cov_{i,j} = \frac{1}{n-1}\sum_{k=1}^{n}(x_{i,k} - mean_i)(x_{j,k} - mean_j),\ i = 1,2,\ldots,m,\quad j = 1,2,\ldots,m. \tag{14}$$

The parameters σ_i, μ_i, $\rho_{i,j}$, i=1,2,…,m, j=1,2,…,m, can be estimated as follows:

$$\sigma_i = \sqrt{\frac{var_i}{\Delta t}},\quad \mu_i = \frac{mean_i}{\Delta t} + \frac{\sigma_i^2}{2},\quad i = 1,2,...,m, \tag{15}$$

$$\rho_{i,j} = \frac{cov_{i,j}}{\sigma_i \sigma_j},\quad i = 1,2,...,m,\quad j = 1,2,...,m. \tag{16}$$

<u>The Heston Model</u>

Let us consider the HM in the case of single assets. According to the HM the stock price volatility is no longer assumed to be constant. Therefore equations (4)-(5) are rewritten as follows:

$$dX(t) = \left(\mu - \frac{\sigma^2}{2}\right)dt + \sigma(t)dW(t),\quad t \geq t_0 \tag{17}$$

$$X(t_0) = 0. \tag{18}$$

Let us define the variance $V(t)$:

$$V(t) = \sigma^2(t),\quad t \geq t_0, \tag{19}$$

The variance $V(t)$ is modelled as a stochastic process satisfying the stochastic differential equation:

$$dV(t) = \gamma(\vartheta - V(t)) + \kappa\sqrt{V(t)}dW^{(1)}(t),\quad t \geq t_0, \tag{20}$$

with initial condition:

$$V(t_0) = v_0. \tag{21}$$

In (20) $\vartheta, \gamma, \kappa$ are positive constant parameters, and $W^{(1)}(t)$ is a standard Wiener process. Let η denote the correlation coefficient between $W(t)$ and $W^{(1)}(t)$, η is assumed to be constant.

The stochastic differential equations (17), (20) with initial conditions (18), (21) constitute the Heston stochastic volatility model for a single asset (Heston, 1993). These equations can be generalized to the case of m assets as follows:

$$dX_i(t) = \left(\mu_i - \frac{V_i(t)}{2}\right)dt + \sqrt{V_i(t)}dW_i(t),\quad t \geq t_0,\ i = 1,2,\ldots,m, \tag{22}$$

$$dV_i(t) = \gamma_i\left(\vartheta_i - V_i(t)\right)dt + \kappa_i\sqrt{V_i(t)}dW_i^{(1)}(t),\ \ t \geq t_0,\ i = 1,2,\ldots,m, \tag{23}$$

$$X_i(t_0) = 0,\ i = 1,2,\ldots,m, \tag{24}$$

$$V_i(t_0) = v_{0i},\ i = 1,2,\ldots,m. \tag{25}$$

In (22), (23) the standard Wiener processes $W_i(t)$ and $W_i^{(1)}(t)$, i=1,2,…,m, j=1,2,…,m, are assumed to be correlated by a constant correlation matrix. In particular let Λ denote the correlation matrix of the 2m-dimensional Wiener process $[W_1(t), W_2(t), \ldots, W_m(t), W_1^{(1)}(t), W_2^{(1)}(t),\ldots, W_m^{(1)}(t)]$, we can represent Λ as the following block matrix:

$$\Lambda = \begin{bmatrix} \Sigma & E \\ E^T & B \end{bmatrix}, \tag{26}$$

where Σ, B are symmetric $m \times m$ matrices and E is a $m \times m$ matrix. We denote with $\rho_{i,j}$ the correlation coefficient between W_i and W_j, with $\beta_{i,j}$ the correlation coefficient between $W_i^{(1)}$ and $W_j^{(1)}$, and with $\eta_{i,j}$ the correlation coefficient between W_i and $W_j^{(1)}$, i=1,2,…,m, j=1,2,…,m. We have:

$$\Lambda = \begin{bmatrix} 1 & \cdots & \rho_{1,m} & \eta_{1,1} & \cdots & \eta_{1,m} \\ \vdots & \ddots & \vdots & \vdots & \ddots & \vdots \\ \rho_{1,m} & \cdots & 1 & \eta_{m,1} & \cdots & \eta_{m,m} \\ \eta_{1,1} & \cdots & \eta_{m,1} & 1 & \cdots & \beta_{1,m} \\ \vdots & \ddots & \vdots & \vdots & \ddots & \vdots \\ \eta_{1,m} & \cdots & \eta_{m,m} & \beta_{1,m} & \cdots & 1 \end{bmatrix} \tag{27}$$

Estimation of the Parameters of the Heston Model

First of all we describe the calibration method used to estimate the parameters of the HM in the case of single assets, i.e. we consider equations (17)-(21).
Let $\tilde{p}(t-t_0, x \mid v_0)$ denote the probability density function of having $X(t)=x$ given $X(t_0)=0$ and $V(t_0)=v_0$. In Dragulescu & Yakovenko, 2002 the following integral representation of $\tilde{p}(t-t_0, x \mid v_0)$ is derived:

$$\tilde{p}(t-t_0, x \mid v_0) = \frac{1}{2\pi}\int_{-\infty}^{+\infty} e^{\Phi(\omega, t-t_0, x, v_0)}\, d\omega, \tag{28}$$

where

$$\Phi(\omega, t-t_0, x, v_0) = i\omega(x + \mu(t-t_0)) - v_0\, \frac{2(\omega^2 - i\omega)}{2\Gamma + \Omega\coth(\Omega)(t-t_0)} - \frac{2\gamma\vartheta}{\kappa^2}\cdot$$

$$\log\left(\cosh\tfrac{\Omega(t-t_0)}{2} + \frac{\Gamma}{\Omega}\sinh\tfrac{\Omega(t-t_0)}{2}\right) + \frac{\gamma\,\Gamma\,\vartheta(t-t_0)}{\kappa^2}, \tag{29}$$

$$\Gamma = \gamma + i\eta\kappa\omega, \tag{30}$$

$$\Omega = \sqrt{\Gamma^2 + \kappa^2(\omega^2 - i\omega)}. \tag{31}$$

Let us consider a set of equally spaced time values t_0, t_1, …,t_n, and let us define $\Delta t = t_k - t_{k-1}$, k=1,2,…,n. Let x_k denote the realized return of a given asset over the time interval $[t_{k-1},t_k]$, k=1,2,…,n. We estimate the parameters $\vartheta, \gamma, \kappa, \eta, \mu$,and the initial variance v_0 using the procedure outlined below:

1. We obtain the empirical probability density function of returns, which we denote with $p_{emp}(\Delta t, x)$, as follows. Let us consider the set of realized returns $\{x_1, x_2, \dots, x_n\}$, and let x_{min} and x_{max} denote the minimum and the maximum of the set $\{x_1, x_2, \dots, x_n\}$ respectively. We divide the interval $[x_{min}, x_{max}]$ into N bins of equal size. Let Δx denote the size of each bin, and let $\bar{x}_k$ denote the center of the k-th bin, k=1,2,…, N. At the centers of the bins, we evaluate the empirical probability density function of returns as follows:

$$p_{emp}(\Delta t, \bar{x}_k) = \frac{n_k}{n}, \quad k = 1,2,\dots, N. \tag{32}$$

2. We treat the initial variance v_0 as a parameter of the HM. More precisely let $p(\Delta t, x)$ denote the probability density function of having $X(t + \Delta t) = x$ given $X(t) = 0$. Instead of formula (28), we consider the following one:

$$p(\Delta t, x) = \frac{1}{2\pi} \int_{-\infty}^{+\infty} e^{\Phi(\omega, \Delta t, x, v_0)} d\omega, \tag{33}$$

that is we assume that the right hand side of equation (28) represents the probability density function of having $X(t + \Delta t) = x$ given $X(t) = 0$ at every time $t \geq t_0$ (and not only at time $t = t_0$ when the variance takes the value v_0).

3. We estimate $\vartheta, \gamma, \kappa, \eta, \mu$,v_0 by minimizing the mean-square deviation, *msd*, between the functions defined in (32) and (33) evaluated at the centers of the bins:

$$msd = \sum_{k=1}^{N} \left[p_{emp}(\Delta t, \bar{x}_k) - p(\Delta t, \bar{x}_k) \right]^2. \tag{34}$$

Note that in order to compute $p(\Delta t, \bar{x}_k)$ the evaluation of the integral appearing in formula (33) is required. In our numerical experience, a fast and accurate numerical approximation of this integral can be computed using Simpson's quadrature rule.

Now let us consider the case of a portfolio of m correlated assets, i.e. equations (22)-(25). We note that the calibration algorithm described above, applied to every single asset of the portfolio, allows to estimate the parameters $\vartheta_i, \gamma_i, \kappa_i, \eta_{i,i}, \mu_i, v_{0i}$, i=1,2,…,m. Therefore the parameters of the HM that still need to be determined are the following elements of the correlation matrix (27): $\rho_{i,j}, \beta_{i,j}, i = 1,2,\dots, m,\ j = 1,2,\dots, m,$ and $\eta_{i,j}, i = 1,2,\dots, m,\ j = 1,2,\dots, m, i \neq j.$

The estimation of these parameters is a very difficult task. In fact, in the case of multiple correlated assets, neither an exact nor an approximate formula for the joint probability distribution of returns of the HM is available in the literature. As a consequence the correlation matrix Λ cannot be estimated by fitting the empirical probability distribution of portfolio returns with some analytical law. Therefore we resort to a heuristic procedure, whose description is given below.

The correlation parameters $\rho_{i,j}, i=1,2,\ldots,m, j=1,2,\ldots,m,$ are evaluated using relations (16), that is the correlations among asset prices are determined as if asset returns were normally distributed. Finally the parameters $\beta_{i,j}, i=1,2,\ldots,m, j=1,2,\ldots,m,$ and $\eta_{i,j}, i=1,2,\ldots,m, \quad j=1,2,\ldots,m, \quad i \neq j,$ are obtained using a matrix completion method that exploits the particular block structure of the matrix (27). A description of this algorithm is presented below.
Let us consider the Cholesky decomposition of the matrix Λ:

$$\Lambda = CC^T, \tag{35}$$

where C is a $2m \times 2m$ lower triangular matrix. Let us denote with $c_{i,j}$ the element of the i-th row and j-th column of the matrix C, $i=1,2,\ldots,2m, \quad j=1,2,\ldots,2m.$ Since Λ has the block structure (26) the elements belonging to the first m rows and the first m columns of matrix C are obtained by Cholesky decomposition of matrix Σ.

The elements of matrix C that still need to be determined, namely $c_{i,k}$, $i=m+1, m+2,\ldots,2m,$ $k=1,2,\ldots,i,$ are obtained using the following relations:

$$c_{i,k} = \begin{cases} \dfrac{\eta_{i-m,i-m}}{c_{i-m,i-m}} \text{ if } \left|\eta_{i-m,i-m}\right| < \left|c_{i-m,i-m}\right|, \quad i=m+1, m+2,\ldots,2m, \quad k=i-m, & \text{(38a)} \\[2em] 0 \text{ if } \left|\eta_{i-m,i-m}\right| < \left|c_{i-m,i-m}\right|, \quad i=m+1, m+2,\ldots,2m, \quad k=1,2,\ldots,i-1, k \neq i-m, & \text{(38b)} \\[2em] \eta_{i-m,i-m}c_{i-m,k} \text{ if } \left|\eta_{i-m,i-m}\right| \geq \left|c_{i-m,i-m}\right|, \quad i=m+1, m+2,\ldots,2m, \quad k=1,2,\ldots,i-1, & \text{(38c)} \\[2em] \sqrt{1-\sum_{j=1}^{i-1} c_{i,j}^2}, \quad i=m+1, m+2,\ldots,2m, \quad k=i. & \text{(38d)} \end{cases}$$

It can be easily checked that, thanks to equations (38a)-(38c) the correlation coefficient between the return and the variance of the i-th asset is equal to $\eta_{i,i}$, $i=1,2,\ldots,m.$ Note also that Equations (38d) ensure that $\beta_{i,i}=1$, $i=1,2,\ldots,m,$ so that the matrix CC^T is a positive definite symmetric matrix with unitary diagonal elements, that is a valid correlation matrix. It is important to observe that, according to equations (38a)-(38c), if $|\eta_{i-m,i-m}| < |c_{i-m,i-m}|$, that is the magnitude of the correlation between the price and the volatility of the i-th asset is smaller than $|c_{i-m,i-m}|$, the i-th row of matrix C has at most two nonzero elements, $i=m+1, m+2,\ldots,2m.$ Moreover we note that if the correlation between the price and the volatility of a given asset is zero, then we would reasonably expect that the correlations of the volatility of that asset with the prices and the volatilities of the other assets are zero as well. This property is respected by the matrix completion algorithm (38a)-(38d). In fact, given an integer i, $i=1,2,\ldots,2m,$ if

$\eta_{i,i} = 0$ from equations (38a), (38b) we obtain $c_{m+i,j} = 0$, $j = 1,2,\ldots,m+i-1$, and hence from relations (27), (35) we have $\eta_{i,k} = 0$, $k = 1,2,\ldots,m$, and $\beta_{i,k} = 0$, $k = 1,2,\ldots,m$, $k \neq i$.

Data

We estimate the parameters of the HM for the following six stocks belonging to the Italian Stock Exchange: Autostrade SpA, Mediobanca SpA, Pirelli & C SpA, RaS Holding, Unicredito Italiano SpA, Snam Rete Gas.

The set of historical data used to derive the empirical probability distribution of returns (32) consists of daily observed asset prices from 17 June 2002 to 15 June 2006. Note that we consider only historical data starting from June 2002 since we want to exclude from our analysis the crash of the Italian financial market due to September 11, 2001. In fact in order to take into account the effects of extreme and unpredictable events such as the the September 11, 2001 terroristic attacks, one should use more ad-hoc models of stock price dynamics, e.g. models with jumps in returns and in volatility (see Eraker, Johannes & Polson, 2003 and references therein). On the other hand we observe that for some of the stocks considered, the historical data are not available on time periods significantly longer than four years (for instance the Snam Rete Gas stock was not quoted on the Italian market before December 2001). We consider only returns on a daily basis, since lower observation frequencies are incompatible with a set of historical data spanning a time horizon of only four years. In fact we have found that the empirical probability distributions of returns computed using time lags longer than 10 days exhibit very irregular shapes. Note also that, since the set of historical data used is relatively small, we do not reject extreme values or bins with low occupation numbers (as done for instance by Daniel, Joseph & Bree, 2005, Dragulescu & Yakovenko, 2002, Prange, Silva & Yakovenko, 2004, Silva & Yakovenko, 2001).

NUMERICAL RESULTS

In this section we present and discuss the numerical results obtained using the calibration method described in Section 4.

The values of the parameters $\vartheta, \gamma, \kappa, \eta, \mu, v_0$ obtained for the six stocks considered are shown in Table 1. Note that in Table 1, as well as in the remainder of the paper, the parameters $\vartheta, \gamma, \kappa, \eta, \mu, v_0$ are expressed in 1/*year* units.

In Figure 1 we compare, for the Pirelli & C. SpA stock, the probability density function of daily returns obtained using the HM (solid line) and the empirical probability density function of returns (dotted line). Note that the probability density function of returns of the HM is computed using formula (28), where the values of $\vartheta, \gamma, \kappa, \eta, \mu, v_0$ are those reported in the third row of Table 1, and $\Delta t = 1\, day$. Figure 1 shows also, for the Pirelli & C. SpA stock, the probability density function of returns obtained using the LM (dashed line), where the drift and volatility parameters are estimated using relations (9).

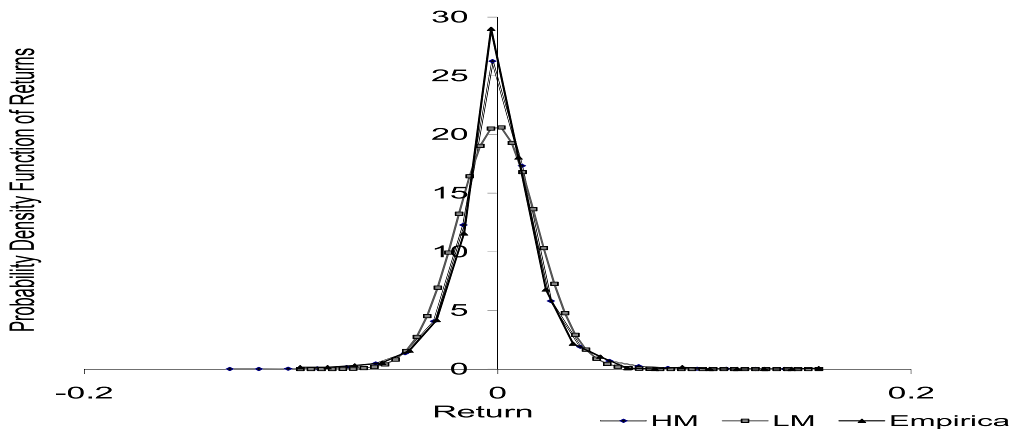
We may note that the HM provides an excellent agreement with historical data and fits the empirical probability distribution of returns far better than the LM. In particular the HM describes the high peaks of the empirical distribution significantly better than the LM. Moreover, although the historical dataset used is probably too small to obtain a sharp description of extreme events, the HM captures the fat tails of the empirical distribution better than the LM. Similar results are obtained also for the other stocks considered. As an example we show in Figure 2 the probability distributions of returns (HM, LM and empirical) obtained for the RaS Holding stock.

Table 1: Parameters of the HM

Number	Name	μ	$v0$	γ	θ	η	κ
1	Autostrade SpA	0.257	0.019	127.6	0.078	0.12	6.70
2	Mediobanca SpA	0.151	0.060	40.30	0.220	0.00	8.30
3	Pirelli & C. SpA	-0.058	0.080	27.5	0.320	0.00	8.90
4	RaS Holding	0.115	0.040	52.1	0.356	-0.07	11.3
5	Unicredito Italiano SpA	0.098	0.012	210.0	0.140	0.04	11.6
6	Snam Rete Gas	0.071	0.030	85.9	0.018	0.03	2.80

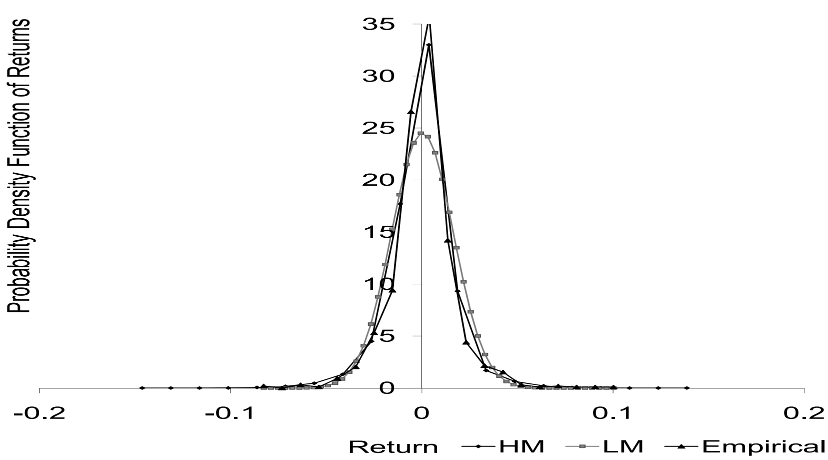
This table shows the values of the parameters of the HM obtained for six stocks traded on the Italian market.

Figure 1: Pirelli & C. SpA, Probability Distributions of Returns



This figure shows the probability density function of daily returns of the Pirelli & C. SpA stock.

Figure 2: RaS Holding, Probability Distributions of Returns



This figure shows the probability density function of daily returns of the RaS Holding stock.

Now let us consider the case of a portfolio composed by the six assets reported in Table 1. The six stocks considered are arranged in vector $[X_1(t), X_2(t), \ldots, X_6(t)]$, whose components appear in equation (22), following the same order as in Table 1. Using the algorithm described in Section 4, we estimate the following correlation matrix:

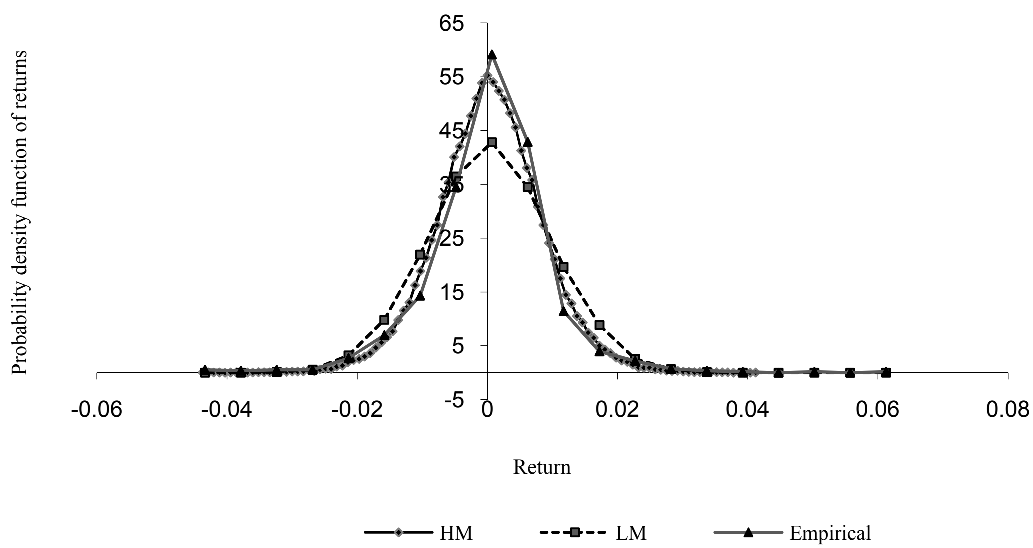
$$\Lambda = \begin{bmatrix} 1 & 0.1971 & 0.1088 & 0.1788 & 0.1590 & 0.1726 & 0.1200 & 0 & 0 & 0 & 0 & 0 \\ 0.1971 & 1 & 0.2758 & 0.4793 & 0.5104 & 0.1137 & 0.0237 & 0 & 0 & 0 & 0 & 0 \\ 0.1088 & 0.2758 & 1 & 0.2216 & 0.2504 & -0.0075 & 0.0131 & 0 & 0 & 0 & 0 & 0 \\ 0.1788 & 0.4793 & 0.2216 & 1 & 0.5333 & 0.1627 & 0.0215 & 0 & 0 & -0.0700 & 0 & 0 \\ 0.1590 & 0.5104 & 0.2504 & 0.5333 & 1 & 0.1082 & 0.0191 & 0 & 0 & -0.0254 & 0 & 0 \\ 0.1726 & 0.1137 & -0.0075 & 0.1627 & 0.1082 & 1 & 0.0207 & 0 & 0 & -0.0092 & 0.0007 & 0.030 \\ 0.1200 & 0.0237 & 0.0131 & 0.0215 & 0.0191 & 0.0207 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & -0.0700 & -0.0254 & -0.0092 & 0 & 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0.0400 & 0.0007 & 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0.0300 & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix}.$$

The initial wealth of the portfolio has been equally distributed among the six stocks, so that every asset gives a relevant contribution to the value of the portfolio. The portfolio return at a given time t is measured as follows:

$$X_P = \log\left(\frac{P_t}{P_{t-1}}\right), \tag{36}$$

In Figure 3 we show the probability distribution of the portfolio returns obtained using the HM, the probability distribution of the portfolio returns obtained using the LM, and the probability distribution of the realized portfolio returns. Note that the probability distribution of both the HM and the LM is computed by Monte Carlo simulation. In particular in the Monte Carlo simulation of the HM, the stochastic differential equations (22), (23) are discretized in time using the Euler-Maruyama scheme (see Kloeden & Platen, 1999). We clearly note that the HM provides a very good description of the realized portfolio returns and fits the empirical distribution of the portfolio returns considerably better than the LM.

Figure 3: Portfolio of Six Stocks, Probability Distributions of Returns



This figure shows the probability density function of daily returns of a portfolio of six stocks traded on the Italian market.

CONCLUSIONS

We have investigated the performance of the Heston stochastic volatility model in describing the probability distribution of returns both in the case of single assets and in the case of asset portfolios. In particular we have proposed a simple method to calibrate the HM based on an integral representation of the exact probability density function of asset returns derived by Dragulescu & Yakovenko (2002). In the case of multiple correlated assets, the correlation parameters are estimated using an ad-hoc matrix completion algorithm.

Using the calibration method proposed in this paper the initial variance of asset returns is treated as an additional parameter of the model. Therefore it is not necessary to assume that the initial variance has stationary probability distribution (as done in previous works), and the value taken by the initial variance can be estimated.

We have used the calibration algorithm presented in this paper to estimate the parameters of the HM for several stocks traded on the Italian financial market. These numerical experiments reveal that, both in the case of single assets and in the case of asset portfolios, our calibration method provides an excellent agreement with historical time series data. Moreover the HM captures the kurtosis effect exhibited by the empirical probability distribution of returns and fits the empirical distribution of returns far better than the LM. Finally we remark that the calibration algorithm proposed in this paper is simple to implement, so we believe that it is very suitable for practical applications.

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IMPLICATIONS OF EUROPEAN TRADING FOR THE NEW YORK STOCK EXCHANGE OPEN

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ABSTRACT

We test the hypothesis that a market maker in New York faces a more competitive market for cross-listed European firms trading simultaneously in their home market during overlapping trading hours as compared to U.S. firms which trade mainly in New York. A sample of seventy two European firms is matched with a control group of U.S. firms, under the same industry and with same liquidity. We find that the mean percentage bid-ask spread for the European firms is significantly smaller than that of the U.S. firms for the opening thirty minutes of trading at the NYSE, even after controlling for liquidity and probability of informed trading. When we compare the percentage bid-ask spreads during the NYSE afternoon after the European markets have closed trading, we find no significant difference. This suggests that the U.S. and the European markets are integrated during the period of overlap and synergies exist between trading venues.

INTRODUCTION

Many European firms cross-list their shares on the U.S. stock exchanges. As of January 27th, 2005 there were 195 companies from 20 different European countries cross-listed on the New York Stock Exchange (NYSE). Most of these European firms cross-list their shares on the NYSE in the form of American Depository Receipts (henceforth ADR). ADR is a derivative instruments backed by home-market ordinary shares. The trend of cross-listing and cross-trading across different equity markets has precipitated a vast body of financial research. The key motivation of most of these studies has been to try and answer the big question: Are the global equity markets integrated or is there evidence of market segmentation? A highly referenced paper in this area has been that of Werner and Kleidon (1996) where they compare a set of U.K. firms cross-listed in the U.S. and a control group of U.K. firms which are not cross-listed in the U.S. Using intraday data they find that, qualitatively, the cross-listed firms do not differ from the locally traded firms in terms of the intraday U shaped price volatility curve- a result one would expect if the markets are segmented.

This result prompted a host of papers amongst which is the one by Lowengrub and Melvin (2000). They use intraday data on a set of German firms and examine the issue of intraday volatility along with volume in a time series setting both before and after the listing date on the U.S. market and find that intraday volatility and volume curves flatten after cross-listing. They conclude that this evidence is consistent with an integrated global trading environment rather than two segmented markets. Eun and Sabherwal (2003) look at the price discovery of Canadian firms on the Toronto Stock Exchange and U.S. exchanges and find that price adjustments due to cross-market information flows take place not only on the U.S. exchange but also on the Toronto Stock Exchange. Grammig, Melvin and Schlag (2004) examine the period of overlap between New York and Germany and find that price discovery for German firms largely occurs in Frankfurt trading. Howe and Ragan (2002) show that the opening volatility of ADRs is lower when the trading of the underlying asset overlaps with the trading of the ADR on the NYSE. In a recent working paper, Moulton and Wei (2004) find that for European firms listed on the NYSE, specialist behavior changes over the day depending upon whether European markets are open or not.

The main idea of this paper is the following. We know that there are many European firms cross-listed as ADRs on the NYSE. When trading opens in New York, for almost two hours there is trading going on simultaneously in the European markets and the NYSE. Table 1 exhibits the trading hour overlap between NYSE and the fifteen major European stock markets.

Table 1: List of European Stock Exchanges

Country	Exchange	Hours	Time ahead of New York	Overlap
Austria	Vienna Stock Exchange	8:30am-5:45pm	6 hrs	2 hrs 15 minutes
Belgium	Euronext Brussels	9 am-5:25 pm	6 hrs	1 hr 55 minutes
Denmark	Copenhagen Stock Exchange	9 am-5 pm	6 hrs	1 hr 30 minutes
Finland	HEX Helsinki	10am-6pm	7 hrs	1 hr 30 minutes
France	Euronext Paris	9 am-5:25 pm	6 hrs	1 hr 55 minutes
Germany	Frankfurt Stock Exchange	9am-8pm	6hrs	4 hrs 30 minutes
Ireland	Irish Stock Exchange	8am-4:30pm	5hrs	2hrs
Italy	Borse Italiana	10am-5:40pm	6 hrs	2 hrs 10 mins
Netherlands	Euronext Amsterdam	9 am-5:25 pm	6 hrs	1 hr 55 minutes
Norway	Oslo Bourse	9am-5pm	6hrs	1hr 30 mins
Portugal	Euronext Lisbon	9 am-5:25 pm	6 hrs	1 hr 55 minutes
Spain	Barcelona Stock Exchange	8:30am-5:45pm	6hrs	2hr 15 mins
Sweden	Stockholm Bourse	9:30am-5:30pm	6hrs	2hrs
Switzerland	Swiss Exchange	9am-5:30pm	6hrs	2hrs
United Kingdom	London Stock Exchange	8am-4:30pm	5hrs	2hrs

So there is a substitute market open for these European stocks in Europe. Now this story should be true for comparable US stocks too (by comparable we mean stocks in same industry with same liquidity). But empirically, we find that US stocks have a clear home bias in terms of trading. US stocks mainly trade at home even thought equal opportunities exist for a European trading. One possible implication, as far as the NYSE market maker (who is dealing with both these stocks) is concerned, is that he faces a more competitive market for the European stocks than the U.S. stocks because of this issue of multimarket trading. The market power of the market maker should, therefore, be reduced for the European firms as compared to the matched U.S. firms. One measure of market power of the market maker in financial markets is the bid-ask spread. So using the simple theoretical background of monopoly versus multimarket trading, we can hypothesize that European stocks will trade at a smaller bid-ask spread than the matched US stocks when markets in the two continents overlap. When trading stops in Europe both sets of stocks should behave same and this difference in bid-ask spread should vanish. We test this theoretical implication. To do that, a sample of seventy two heavily traded European firms is collected and matched with a group of U.S. firms on the basis of industry and liquidity. Table 2 documents the sample of U.S and European firms used in this study.

We then study the high frequency bid and ask quotes for two time periods: the first thirty minutes of trading, from 9:30-10 am, when the trading hours in the NYSE and the European markets overlap and from 2:30-3 pm when only the NYSE is trading and all the European markets have closed. We use high frequency tick-by-tick data from the TAQ (Trade and Quote) database and three months of data from September-November, 2000 and compare the percentage bid-ask spreads between the European and the U.S. firms. Our hypothesis is that, because of competition from overseas home markets during the NYSE morning, the European firms should trade at a smaller bid-ask spread than the U.S. firms. But this difference should vanish during the NYSE afternoon when all the European markets have closed.

Of course bid-ask spreads might also be driven by liquidity and informed trading in a stock. The idea of bid-ask spreads being driven by informed trading follows from the theory that the risk-averse market maker will set bigger bid-ask spreads to compensate for the risk exposure when there is a higher probability of trading with a privately informed trader. This would be especially true during the NYSE morning when we would expect privately informed traders to be more active. To account for that, we estimate the probability of informed trading (PIN) in our sample using the method of Easley, O'Hara, Kiefer and Paperman (1996).

Table 2: Sample of U.S. and European Firms

European Firm	Ticker	U.S. Firm	Ticker	Industry
Publicis Groupe S.A.	PUB	Harte Hanks.	HHS	Advertising
Autoliv Inc.	ALV	Tower Automotive.	TWR	Autoparts
Banco Bilbao Viscaya Argentaria.	BBV	Bancorp South.	BXS	Banking
Banco Santander Central Hispanio S.A.	STD	Bayview Capital.	BVC	Banking
ABN AMRO Bank.	ABN	Cullen/Frost Bankers.	CFR	Banking
Allied Irish Banks.	AIB	First Fed Financial Corp.	FED	Banking
Barclays Plc.	BCS	Valley National Bancorp.	VLY	Banking
Credit Suisse Group.	CSR	Bankatlantic Bancorp.	BBX	Banking
Deutsche Bank A.G.	DB	Chittenden Corporation.	CZN	Banking
HSBC Holdings.	HBC	M&T Bancorp.	MTB	Banking
Sanpaolo IMI.	IMI	Community Bank System.	CBU	Banking
UBS A.G.	UBS	Commercial Federal Corp.	CFB	Banking
Lloyds TSB Group Plc.	LYG	First Commonwealth.	FCF	Banking
Serono S.A.	SRA	Theragenic Corp.	TGX	Biotechnology
Vivendi Universal.	V	Hearst Arghyle Television.	HTV	Broadcasting
Hanson Plc.	HAN	Ameron International.	AMN	Building materials
Luxottica Group.	LUX	Guess Inc et al.	GES	Clothes and Fabrics
Alcatel.	ALA	American Tower Corp.	AMT	Communications Technology
Siemens A.G.	SI	Cable design Corp.	CDT	Communications Technology
Nokia Corporation.	NOK	Corning Inc.	GLW	Communications Technology
BASF A.G.	BF	Spartech Corp.	SEH	Commodity Chemicals
Bayer A.G.	BAY	Wellman Inc.	WLM	Commodity Chemicals
Celanese A.G.	CZ	NL Industries.	NL	Commodity Chemicals
Royal Phillips Electronics.	PHG	Harman Intl.	HAR	Consumer Electronics
Diageo Plc.	DEO	Brown Forman.	BFB	Distillers and Brewers
Endesa S.A.	ELE	Unisource Energy.	UNS	Electric Utilities
E.ON G.	EON	CH Energy.	CHG	Electric Utilities
Scottish Power UK Plc.	SPI	El Paso Electric.	EE	Electric Utilities
Cable and Wireless.	CWP	IDT Corporation.	IDT	Fixed line Communications
Deutsche Telekom A.G.	DT	Cincinnati Bell.	CBB	Fixed line Communications
France Telecom.	FTE	Sprint Corporation.	SDE	Fixed line Communications
TDC A/S.	TLD	Centurytel.	CTLPRA	Fixed line Communications
Telefonica S..A.	TEF	BCE Inc.	BCE	Fixed line Communications
Groupe Danone.	DA	M&F Worldwide.	MFW	Food Products
Cadbury Schweppes Plc	CSG	Ralcorp Holdings.	RAH	Food Products
Unilever N.V.	UN	Mccormick &Co.	MKC	Food Products
Delhaize Group	DEG	Smart and Final.	SMF	Food Retailers &Wholesellers
Natuzzi SPA.	NTZ	Fedders Corp.	FJC	Furnishing and Appliance
Royal Ahold.	AHO	Winn Dixie Stores.	WIN	Food Retailers &Wholesellers
Aegon N.V.	AEG	CNA Financial.	CNA	Full line Insurance
Allianz A.G.	AZ	Horace Mann Educators.	HMN	Full line Insurance

European Firm	Ticker	U.S. Firm	Ticker	Industry
AXA.	AXA	Stancorp Financial.	SFG	Full line Insurance
Royal and Sun Alliance Insurance Grp Plc	RSA	FBL Financial Grp.	FFG	Full line Insurance
Chicago Bridge & Iron Co.	CBI	MasTec Inc.	MTZ	Heavy Construction
Adecca S.A.	ADO	Crawford and Company.	CRDB	Industrial Services
AMVESCAP Plc.	AVZ	Gabelli Asset Mgt.	GBL	Investment Services
ING Group.	ING	Nationwide Fin Services.	NFS	Life Insurance
BP Plc.	BP	ConocoPhillips.	COP	Major Oil Companies
Royal Dutch Petroleum Co.	RD	Marathon Oil Corp.	MRO	Major Oil Companies
TOTAL S.A.	TOT	Unocal.	UCL	Major Oil Companies
Alcon Inc.	ACL	Apogent Technology.	AOT	Medical Supplies
Rio Tinto.	RTP	Cleveland Cliffs.	CLF	Mining
Core labs.	CLB	Carbo Ceramics.	CRR	Oil Drilling
Stora Enso.	SEO	Buckeye Tech.	BKI	Paper Products
UPM-Kymmene Corporation.	UPM	Schweitzer Mauduit Intl.	SWM	Paper products
Aventis S.A.	AVE	Bradley Pharmaceuticals.	BDY	Pharmaceuticals
GlaxoSmithKline Plc.	GSK	Pharmaceutical Resources	PRX	Pharmaceuticals
Novartis.	NVS	Medicis Pharmaceuticals.	MRX	Pharmaceuticals
AstraZeneca Grp.	AZN	Alpharma Inc.	ALO	Pharmaceuticals
Elan Corp.	ELN	Mylan Labs.	MYL	Pharmaceuticals
Schering A.G.	SHR	KV Pharmaceauticals.	KVB	Pharmaceuticals
Wilis Grp.	WSH	Allmerica Financial Corp.	AFC	Property and Casualty Insurance
Pearson Plc.	PSO	Coachman Industries.	COA	Recreational Products and Services
Carnival Plc.	CUK	Dover Motorsports.	DVD	Recreational Products and Services
Infineon Technologies	IFX	Memc Electronic Materials.	WFR	Semiconductors
ST Microelectronics N.V.	STM	Fairchild Semiconductors.	FCS	Semiconductors
Imperial Chemical Industries Plc.	ICI	Arch Chemicals.	ARJ	Speciality Chemicals
Syngenta.	SYT	Rogers Corp.	ROG	Speciality Chemicals
SAP A.G.	SAP	Cadence Design System.	CDN	Technology,Software
Gallagher Group Plc.	GLH	Standard Commercial Corp	STW	Tobacco

The first column of the table reports the list of European firms, the second column their NYSE ticker symbol, the third column the matching U.S. firms. The fifth column lists the name of the industry to which the pair of European and U.S. firm in that row belongs. We have used the subgroup classifications under the Dow Jones Global Classification Standard.

As for liquidity, the more liquid the trading in a stock the smaller the bid-ask spread. We use the consolidated number of trades in a stock as a measure of liquidity. We then compare the percentage bid-ask spreads between the European and the U.S. stocks for the two different time periods of the day, after controlling for the effect of liquidity and the extent of informed trading. Our results show that the European firms trade at a significantly smaller bid-ask spread during the NYSE morning period. During the NYSE afternoon, however, the differences in bid-ask spreads vanish. This indicates that our initial hypothesis is true, the NYSE market maker does face competition from European trading during the NYSE morning. This also indicates that the U.S. and the European markets are integrated during the morning period of overlap.

The paper proceeds as follows. First we discuss the theory of informed trading using an earlier paper by Easley, Kiefer, O'Hara and Paperman, 1996 (henceforth EKOP). Next we discuss the empirical evidence on the bid-ask spread is presented and finally we conclude.

INFORMED TRADING

The bid-ask spreads may also be driven also by the extent of informed trading in a stock. Risk-averse market makers tend to set bigger bid-ask spreads to compensate for the exposure to privately informed traders. So when we test our hypothesis if European stocks trade at a smaller bid-ask spread than U.S. stocks because of market competitiveness, we want to make sure we control for the effect of informed trading, if any. This effect is especially important during the NYSE morning when informed traders are expected to quickly trade on their private information. It is difficult to estimate if there is private information based trading going on in a stock. EKOP (1996) have developed an empirical technique to test for the presence of informed trading in a stock. The idea is to use the information in the trade data to estimate the probability of informed trading. Specifically, they use a continuous time sequential model and develop a likelihood function to use in the estimation. The setup of their model is as follows:

- One risk neutral market maker and many informed and uninformed traders.
- Individuals trade a single risky asset and money with a market maker over $i = 1,.....I$ days. Within each trading day time is continuous and is indexed by $t \in [0,T]$.
- Prior to the beginning of each trading day, nature determines whether an information event happens. Information events are independent and occur with probability d. These events are bad news with probability e and good news with probability $1-e$.
- $(V_i)_{i=1}^{I}$ are the random variables that give the value of the asset at the end of day I.
- Uninformed buyer and seller order arrivals are Poisson processes and the rate of arrival per minute is κ. Informed buyer and seller order arrivals are also Poisson and the rate of arrival per minute is ϖ. Order imbalance is expected to occur with informed trader activity.
- If a privately informed trader observes a bad signal he sells, if he observes a good signal he buys.
- The market maker is a Bayesian and he updates his belief about an information event by looking at the arrival of trade and rate of trading.

EKOP derives the probability of informed trading as $PIN = \dfrac{d\varpi}{d\varpi + 2\kappa}$.

They use a structural model to estimate the parameters d, ϖ, κ. The likelihood function is derived as the following:

$$L(B,S \mid \mathrm{H}) = (1-d) * e^{-\kappa T} \frac{(\kappa T)^B}{B!} e^{-\kappa T} \frac{(\kappa T)^S}{S!} + de * e^{-\kappa T} \frac{(\kappa T)^B}{B!} e^{-(\varpi+\kappa)T} \frac{[(\varpi+\kappa)T]^S}{S!}$$
$$+d(1-e) * e^{-\kappa T} \frac{(\kappa T)^S}{S!} e^{-(\varpi+\kappa)T} \frac{[(\varpi+\kappa)T]^B}{B!} \qquad (1)$$

Where $B =$ number of buys in a day, $S =$ number of sells in a day, H is the parameter vector. EKOP use the Lee and Ready [1990] algorithm to classify each trade as a buy or sell. The likelihood of observing data $M = (B_i, S_i)_{i=1}^{I}$ over I days is just the product of the daily likelihoods,

$$L(M \mid \mathrm{H}) = \prod_{i=1}^{I} L(\mathrm{H} \mid B_i, S_i).$$

For our paper we use this technique to measure the probability of informed trading for our sample of European and U.S. stocks. We use the Lee and Ready technique to classify each trade as buy or sell. Then we maximize the likelihood function and find the parameter estimates and obtain the PIN value for each of our stocks. Each PIN value is a number between 0 and 1. Tables 3A and 3B document the PIN for the entire sample for the two time periods 9:30-10am and 2:30-3 pm respectively.

Table 3A: Probability of Informed Trading for 9:30-10 AM

US Firm	PIN	European Firm	PIN
Gerber Scientific	0.428	ABB Limited	0.324
Graco Inc	0.219	Mettler Toledo	0.205
Harte Hanks	0.265	Publicis Groupe S.A.	0.785
Tower Automotive	0.274	Autoliv Inc.	0.254
Bancorp South	0.257	Banco Bilbao	0.228
Bayview Capital	0.401	Banco Santander Central	0.311
Cullen/Frost Bankers	0.189	ABN AMRO Bank	0.144
First Fed Financial Corp	0.295	Allied Irish Banks	0.226
Valley National Bancorp	0.209	Barclays Plc.	0.225
Bankatlantaic Bancorp	0.246	Credit Suisse	0.231
Chittenden Corporation	0.195	Deutsche Bank	0.209
M&T Bancorp	0.138	HSBC holdings	0.171
Community Bank System	0.389	Sanpaolo IMI	0.265
Commercial Federal Corp	0.253	UBS AG	0.18
First Commonwealth	0.523	Lloyds TSB Group	0.202
Theragenic Corp	0.301	Serono	0.233
Hearst Arghyle Television	0.282	Vivendi Universal	0.186
Ameron International	0.33	Hanson Plc.	0.471
Guess Inc et al	0.345	Luxottica	0.302
American Tower Corp	0.171	Alcatel	0.125
Cable design Corp	0.278	Siemens AG	0.128
Corning Inc	0.159	Nokia Corporation	0.083
Spartech Corp	0.34	BASF AG	0.186
Wellman Inc	0.251	Bayer AG	0.178
NL Industries	0.335	Celanese AG	0.28
Harman Intl	0.193	Royal Phillips	0.144
Brown Forman	0.278	DIAGEO plc	0.13
Unisource Energy	0.197	Endesa SA	0.288
CH Energy	0.43	E ON G	0.279
El Paso Electric	0.368	Scottish Power UK plc	0.286
IDT Corporation	0.288	Cable and Wireless	0.279
Cincinnati Bell	0.273	Deutsche Telekom	0.188
Sprint Corporation	0.21	France Telecom	0.232
Centurytel	0.595	TDC A/S	0.843
BCE Inc	0.181	Telefonica	0.16
M&F Worldwide	0.317	Groupe Danone	0.111
Ralcorp Holdings	0.266	Cadbury Schweppes	0.196
Mccormick &Co	0.193	Unilever	0.226
Smart and Final	0.34	Delhaize Group	0.293
Fedders Corp	0.363	Natuzzi SPA	0.205
Winn Dixie Stores	0.155	Royal Ahold	0.155
CNA Financial	0.155	AEGON	0.149
Horace Mann Educators	0.3	Allianz	0.222
Stancorp Financial	0.236	AXA	0.193
FBL Financial Grp	0.18	Royal and Sun Alliance	0.263
MasTec Inc	0.26	Chicago Bridge&Iron	0.221
Crawford and Company	0.693	Adecca	0.423
Gabelli Asset Mgt	0.336	AMVESCAP Plc	0.193
Nationwide Fin Services	0.126	ING Group	0.191
ConocoPhillips	0.131	BP Plc.	0.114
Marathon Oil Corp	0.125	Royal Dutch Petroleum	0.168
Unocal	0.141	TOTAL S.A.	0.193
Apogent Technology	0.226	Alcon Inc	0.213

US Firm	PIN	European Firm	PIN
Cleveland Cliffs	0.251	Rio Tinto	0.22
Carbo Ceramics	0.26	Core labs	0.319
Buckeye Tech	0.348	Stora Enso	0.216
Schweitzer Mauduit Intl	0.373	UPM-Kymmene Corporation	0.518
Bradley Pharmaceuticals	0.226	Aventis	0.15
Pharmaceutical Resources	0.137	GlaxoSmithKline plc	0.197
Medicis Pharmaceuticals	0.277	Novartis	0.109
Alpharma Inc	0.114	AstraZeneca Grp	0.14
Mylan Labs	0.152	Elan Corp	0.162
KV Pharmaceauticals	0.578	Schering Aktiengesellschaft	0.297
Allmerica Financial Corp	0.22	Wilis Grp.	0.205
Coachman Industries	0.281	Pearson plc.	0.245
Dover Motorsports	0.535	Carnival Plc.	0.286
Memc Electronic Materials	0.209	Infineon Technologies	0.086
Fairchild Semiconductors	0.133	STMicroelectronics	0.647
Arch Chemicals	0.376	Imperial Chemical Industries PLC	0.245
Rogers Corp	0.374	Syngenta	0.365
Cadence Design System	0.1	SAP	0.133
Standard Commercial Corp	0.376	Gallagher Group Plc	0.359

The first column of the table reports the list of U.S. firms, the second column their value of probability of informed trading (PIN). The fifth and sixth column reports the matching European firms and the value of probability of informed trading (PIN).

Table 3B: Probability of Informed Trading for 2:30-3 PM

U.S. Firm	PIN	European Firm	PIN
Gerber Scientific	0.414	ABB Limited	0.239
Graco Inc	0.217	Mettler Toledo	0.259
Harte Hanks	0.286	Publicis Groupe S.A.	1.000
Tower Automotive	0.288	Autoliv Inc.	0.184
Bancorp South	0.303	Banco Bilbao	0.242
Bayview Capital	0.432	Banco Santander Central	0.258
Cullen/Frost Bankers	0.185	ABN AMRO Bank	0.070
First Fed Financial Corp	0.302	Allied Irish Banks	0.305
Valley National Bancorp	0.179	Barclays Plc.	0.199
Bankatlantic Bancorp	0.195	Credit Suisse	0.209
Chittenden Corporation	0.199	Deutsche Bank	0.277
M&T Bancorp	0.197	HSBC holdings	0.190
Community Bank System	0.185	Sanpaolo IMI	0.095
Commercial Federal Corp	0.227	UBS AG	0.228
First Commonwealth	0.290	Lloyds TSB Group	0.293
Theragenic Corp	0.280	Serono	0.336
Hearst Arghyle Television	0.254	Vivendi Universal	0.188
Ameron International	0.152	Hanson Plc.	0.185
Guess Inc et al	0.292	Luxottica	0.290
American Tower Corp	0.194	Alcatel	0.176
Cable design Corp	0.243	Siemens AG	0.234
Corning Inc	0.116	Nokia Corporation	0.088
Spartech Corp	0.211	BASF AG	0.193
Wellman Inc	0.252	Bayer AG	0.280
NL Industries	0.191	Celanese AG	0.260
Harman Intl	0.148	Royal Phillips	0.123
Brown Forman	0.168	DIAGEO plc	0.379
Unisource Energy	0.263	Endesa SA	0.305

U.S. Firm	PIN	European Firm	PIN
CH Energy	0.296	E ON G	0.144
El Paso Electric	0.219	Scottish Power UK plc	0.137
IDT Corporation	0.327	Cable and Wireless	0.395
Cincinnati Bell	0.240	Deutsche Telekom	0.175
Sprint Corporation	0.437	France Telecom	0.315
Centurytel	0.576	TDC A/S	0.750
BCE Inc	0.251	Telefonica	0.278
M&F Worldwide	0.355	Groupe Danone	0.273
Ralcorp Holdings	0.127	Cadbury Schweppes	0.208
Mccormick &Co	0.249	Unilever	0.265
Smart and Final	0.310	Delhaize Group	0.626
Fedders Corp	0.312	Natuzzi SPA	0.338
Winn Dixie Stores	0.190	Royal Ahold	0.207
CNA Financial	0.212	AEGON	0.177
Horace Mann Educators	0.251	Allianz	0.318
Stancorp Financial	0.175	AXA	0.202
FBL Financial Grp	0.430	Royal and Sun Alliance	0.466
MasTec Inc	0.242	Chicago Bridge&Iron	0.256
Crawford and Company	0.628	Adecca	0.471
Gabelli Asset Mgt	0.274	AMVESCAP Plc	0.277
Nationwide Fin Services	0.113	ING Group	0.175
ConocoPhillips	0.105	BP Plc.	0.183
Marathon Oil Corp	0.160	Royal Dutch Petroleum	0.186
Unocal	0.175	TOTAL S.A.	0.183
Apogent Technology	0.252	Alcon Inc	0.229
Cleveland Cliffs	0.253	Rio Tinto	0.261
Carbo Ceramics	0.210	Core labs	0.303
Buckeye Tech	0.318	Stora Enso	0.192
Schweitzer Mauduit Intl	0.296	UPM-Kymmene Corporation	0.180
Bradley Pharmaceuticals	0.222	Aventis	0.191
Pharmaceutical Resources	0.182	GlaxoSmithKline plc	0.183
Medicis Pharmaceuticals	0.184	Novartis	0.187
Alpharma Inc	0.222	AstraZeneca Grp	0.149
Mylan Labs	0.131	Elan Corp	0.144
KV Pharmaceauticals	0.664	Schering Aktiengesellschaft	0.652
Allmerica Financial Corp	0.179	Wilis Grp.	0.202
Coachman Industries	0.387	Pearson plc.	0.286
Dover Motorsports	0.411	Carnival Plc.	0.252
Memc Electronic Materials	0.208	Infineon Technologies	0.238
Fairchild Semiconductors	0.153	STMicroelectronics	0.190
Arch Chemicals	0.338	Imperial Chemical Industries PLC	0.273
Rogers Corp	0.265	Syngenta	0.227
Cadence Design System	0.177	SAP	0.166
Standard Commercial Corp	0.449	Gallagher Group Plc	0.265

The first column of the table reports the list of U.S. firms, the second column their value of probability of informed trading (PIN). The fifth and sixth columns report the European firms and the value of probability of informed trading (PIN)

EMPIRICAL EVIDENCE

For our purposes we needed European stocks which are highly liquid and have active trading going on in their home markets in Europe when their ADRs start to trade at the NYSE. The stocks were selected from those European countries whose major stock markets have substantial overlapping trading hours with the NYSE. The major stock exchanges selected were those of the following 15 European countries: Austria, Belgium, Denmark, Finland, France, Germany, Italy, Ireland, Netherlands, Norway, Portugal, Sweden, Switzerland, Spain and U.K. The sample consists of seventy two heavily traded European stocks from the above list of countries and a control group of U.S. stocks, matched by the consolidated number of trades in 2003 under the same industry. The NYSE uses the Dow Jones Global Classification Standard which divides the firms into 10 economic sectors, 18 market sectors, 51 industry groups and 89 subgroups. Each of the subgroups was examined to pick the European firm and the matched U.S. firm. Table 2 lists the firms. High frequency tick-by-tick bid and offer quotes are used from the NYSE Trade and Quote(TAQ) database which consists of time stamped intraday transactions data for all securities listed on the NYSE and American Stock Exchange (AMEX) as well as NASDAQ National Market System (NMS) and Small Cap issues. The bid and offer quotes and the trades were extracted for the sample for the months of September, October and November, 2003 for the two time periods, 9:30-10 am and 2:30-3 pm. There are 63 trading days in the sample.

Evidence on Bid-ask Spread

The data were sorted and stacked according to each firm, date, time and a variable that identifies each minute of the trading day. Then the percentage bid-ask spread, *perspread*, where $perspread = \frac{Spread}{Midprice}$ is computed for each quote in the dataset. For each firm the mean *perspread* for the entire trading period is computed. So we obtain one mean *perspread* value for each of our firms in both the samples. Table 4A lists the *perspread* values for the sample for the 9:30-10 am period and Table 4B lists the per spread values for the 2:30-3 pm period.

Sample Period 9:30-10 AM

We expected that within each set of U.S. and European firms the mean perspread should go down from the less liquid to the more liquid stocks. This is compatible with the intuition that the higher the liquidity, the lower should be the transactions cost, bid-ask spread being a widely used measure of transaction cost. We computed the correlation coefficient between the mean *perspread* and the number of trades for each of the set of European and U.S. firms and found that the correlation coefficient is -0.51771 for the U.S. firms and -0.72960 for the European firms. So the initial evidence suggests that more liquid the stock, smaller the bid-ask spread would be. Focusing on the morning sample, it was also found that the mean *perspread* for the U.S sample is 0.00522027 and the mean *perspread* for the European sample is 0.002976261. So the mean *perspread* for the European sample is smaller than that of the U.S. sample by approximately 23 basis points. The associated P value is 0. Out of each of the pairs of firms it was found that for 17 pairs of firms the European firm has a bigger *perspread* and for the rest of the 55 pairs the U.S. firm has a bigger *perspread*. So overwhelmingly the U.S. firms trade at a bigger bid-ask spread than the European firms.

Table 4 A: Percentage Bid-ask spread Analysis for 9:30-10 AM

U.S. Firm	Perspread	European Firm	Perspread	X
Gerber Scientific	0.0112014	ABB Limited	0.0066126	US
Graco Inc	0.0017825	Mettler Toledo	0.0023529	US
Harte Hanks	0.0034331	Publicis Groupe S.A.	0.0045	EUR
Tower Automotive	0.0108557	Autoliv Inc.	0.001491	EUR
Bancorp South	0.0033892	Banco Bilbao	0.00409	US
Bayview Capital	0.0038202	Banco Santander Central	0.0043657	EUR
Cullen/Frost Bankers	0.0017393	ABN AMRO Bank	0.0015749	EUR
First Fed Financial Corp	0.0024	Allied Irish Banks	0.003946	US
Valley National Bancorp	0.0025414	Barclays Plc.	0.0019518	EUR

U.S. Firm	Perspread	European Firm	Perspread	X
Bankatlantaic Bancorp	0.0046525	Credit Suisse	0.0017884	US
Chittenden Corporation	0.022499	Deutsche Bank	0.00155	US
M&T Bancorp	0.001163	HSBC holdings	0.00063502	US
Community Bank System	0.0046619	Sanpaolo IMI	0.0036868	US
Commercial Federal Corp	0.0020973	UBS AG	0.001186	US
First Commonwealth	0.006732	Lloyds TSB Group	0.0020292	US
Theragenic Corp	0.0115576	Serono	0.0032615	US
Hearst Arghyle Television	0.0021103	Vivendi Universal	0.0018581	US
Ameron International	0.0046619	Hanson Plc.	0.0044619	US
Guess Inc et al	0.0092812	Luxottica	0.0038163	US
American Tower Corp	0.0037735	Alcatel	0.0020248	US
Cable design Corp	0.0049271	Siemens AG	0.0013626	US
Corning Inc	0.0016237	Nokia Corporation	0.00098545	US
Spartech Corp	0.0053746	BASF AG	0.0018222	US
Wellman Inc	0.006196	Bayer AG	0.0021906	US
NL Industries	0.0071415	Celanese AG	0.004662	US
Harman Intl	0.0016568	Royal Phillips	0.0012552	US
Brown Forman	0.0011459	DIAGEO plc	0.00086415	US
Unisource Energy	0.0040211	Endesa SA	0.0034832	US
CH Energy	0.0047346	E ON G	0.0022141	US
El Paso Electric	0.0050564	Scottish Power UK plc	0.0022458	US
IDT Corporation	0.0031806	Cable and Wireless	0.0060163	US
Cincinnati Bell	0.0053963	Deutsche Telekom	0.0020331	EUR
Sprint Corporation	0.0111713	France Telecom	0.002422	US
Centurytel	0.0080831	TDC A/S	0.0088314	US
BCE Inc	0.0016433	Telefonica	0.0018127	EUR
M&F Worldwide	0.0091046	Groupe Danone	0.0021907	EUR
Ralcorp Holdings	0.0037014	Cadbury Schweppes	0.0018732	US
Mccormick &Co	0.001295	Unilever	0.00084575	US
Smart and Final	0.0142741	Delhaize Group	0.0041157	US
Fedders Corp	0.0142426	Natuzzi SPA	0.0073232	US
Winn Dixie Stores	0.0025971	Royal Ahold	0.0029	US
CNA Financial	0.0026223	AEGON	0.0020992	EUR
Horace Mann Educators	0.004842	Allianz	0.0036317	US
Stancorp Financial	0.0018497	AXA	0.0020504	US
FBL Financial Grp	0.0077774	Royal and Sun Alliance	0.0125151	EUR
MasTec Inc	0.0042873	Chicago Bridge&Iron	0.0040511	EUR
Crawford and Company	0.0167837	Adecca	0.0058341	US
Gabelli Asset Mgt	0.0032635	AMVESCAP Plc	0.0041915	US
Nationwide Fin Services	0.0032628	ING Group	0.0018504	EUR
ConocoPhillips	0.00065312	BP Plc.	0.00052596	US
Marathon Oil Corp	0.0011581	Royal Dutch Petroleum	0.00048054	US
Unocal	0.0014577	TOTAL S.A.	0.000797	US
Apogent Technology	0.002149	Alcon Inc	0.0014776	US
Cleveland Cliffs	0.0052459	Rio Tinto	0.0013166	US
Carbo Ceramics	0.0025212	Core labs	0.005312	US
Buckeye Tech	0.0101558	Stora Enso	0.0034648	EUR
Schweitzer Mauduit Intl	0.0030526	UPM-Kymmene Corporation	0.0049122	US
Bradley Pharmaceuticals	0.0043029	Aventis	0.0012985	EUR
Pharmaceutical Resources	0.0017099	GlaxoSmithKline plc	0.00082006	US
Medicis Pharmaceuticals	0.001419	Novartis	0.00085809	US
Alpharma Inc	0.002843	AstraZeneca Grp	0.000152	US

U.S. Firm	Perspread	European Firm	Perspread	X
Mylan Labs	0.0011831	Elan Corp	0.00587	US
KV Pharmaceauticals	0.0067368	Schering Aktiengesellschaft	0.0030094	EUR
Allmerica Financial Corp	0.0017962	Wilis Grp.	0.0019741	US
Coachman Industries	0.0102348	Pearson plc.	0.0055529	EUR
Dover Motorsports	0.0191393	Carnival Plc.	0.0034355	US
Memc Electronic Materials	0.002989	Infineon Technologies	0.0019379	US
Fairchild Semiconductors	0.0015876	STMicroelectronics	0.00090429	US
Arch Chemicals	0.005199	Imperial Chemical Industries PLC	0.004133	US
Rogers Corp	0.0033578	Syngenta	0.0050665	US
Cadence Design System	0.00141	SAP	0.00087785	EUR
Standard Commercial Corp	0.0095623	Gallagher Group Plc	0.0026384	US

The first column of the table reports the list of U.S firms, the second column the NYSE ticker symbols and the third column the mean percentage bid-ask spread for the trading period. The columns four, five and six report the same for the European firms and the seventh column reports the variable "X". If X= "US" ("EUR") it means the U.S.(European) firm in that pair has a higher percentage bid-ask spread. The column header perspread denotes the percentage bid-ask spread.

Table 4B: Percentage Bid-ask Spread Analysis for 2:30-3:00 PM

U.S. Firm	Perspread	European Firm	Perspread	X
Gerber Scientific	0.0010091	ABB Limited	0.008576465	EUR
Graco Inc	0.0012433	Mettler Toledo	0.001591212	US
Harte Hanks	0.0045953	Publicis Groupe S.A.	0.000873659	EUR
Tower Automotive	0.0019991	Autoliv Inc.	0.006547056	EUR
Bancorp South	0.0008353	Banco Bilbao	0.002316949	US
Bayview Capital	0.0031356	Banco Santander Central	0.003066669	EUR
Cullen/Frost Bankers	0.0024092	ABN AMRO Bank	0.002789415	EUR
First Fed Financial Corp	0.0013428	Allied Irish Banks	0.001775627	US
Valley National Bancorp	0.002961	Barclays Plc.	0.000840659	US
Bankatlantaic Bancorp	0.0031099	Credit Suisse	0.001094616	US
Chittenden Corporation	0.006	Deutsche Bank	0.00408147	EUR
M&T Bancorp	0.0026336	HSBC holdings	0.002664052	EUR
Community Bank System	0.002539	Sanpaolo IMI	0.004432886	US
Commercial Federal Corp	0.0028195	UBS AG	0.0005872	US
First Commonwealth	0.0030552	Lloyds TSB Group	0.002749376	EUR
Theragenic Corp	0.0010034	Serono	0.004124914	EUR
Hearst Arghyle Television	0.0026937	Vivendi Universal	0.002755641	EUR
Ameron International	0.0013878	Hanson Plc.	0.002380152	US
Guess Inc et al	0.0008868	Luxottica	0.000471316	US
American Tower Corp	0.0028004	Alcatel	0.002644023	US
Cable design Corp	0.0030788	Siemens AG	0.003014056	EUR
Corning Inc	0.0012939	Nokia Corporation	0.00242495	US
Spartech Corp	0.006263	BASF AG	0.001501128	EUR
Wellman Inc	0.00042	Bayer AG	0.00325786	US
NL Industries	0.0113109	Celanese AG	0.005144409	EUR
Harman Intl	0.0018568	Royal Phillips	0.004104	US
Brown Forman	0.0065276	DIAGEO plc	0.00300783	EUR
Unisource Energy	0.0011638	Endesa SA	0.002007695	US
CH Energy	0.0149562	E ON G	0.005626067	US
El Paso Electric	0.0030378	Scottish Power UK plc	0.000626437	US
IDT Corporation	0.0044605	Cable and Wireless	0.00173422	EUR
Cincinnati Bell	0.0010326	Deutsche Telekom	0.004545848	EUR
Sprint Corporation	0.0015501	France Telecom	0.009541478	US
Centurytel	0.0046059	TDC A/S	0.004390942	US
BCE Inc	0.0115923	Telefonica	0.003618708	EUR
M&F Worldwide	0.0023867	Groupe Danone	0.002643699	US

U.S. Firm	Perspread	European Firm	Perspread	X
Ralcorp Holdings	0.005692	Cadbury Schweppes	0.000763981	EUR
Mccormick &Co	0.0010919	Unilever	0.004277899	US
Smart and Final	0.0018762	Delhaize Group	0.000598956	US
Fedders Corp	0.0092497	Natuzzi SPA	0.005679647	EUR
Winn Dixie Stores	0.0008523	Royal Ahold	0.002159306	EUR
CNA Financial	0.0013583	AEGON	0.004333959	EUR
Horace Mann Educators	0.0020099	Allianz	0.002120634	EUR
Stancorp Financial	0.0013177	AXA	0.002666733	US
FBL Financial Grp	0.0020905	Royal and Sun Alliance	0.002047568	US
MasTec Inc	0.0060751	Chicago Bridge&Iron	0.00115963	US
Crawford and Company	0.0062644	Adecca	0.001226834	EUR
Gabelli Asset Mgt	0.0005691	AMVESCAP Plc	0.005081196	EUR
Nationwide Fin Services	0.0005573	ING Group	0.000673902	US
ConocoPhillips	0.0011367	BP Plc.	0.001036373	EUR
Marathon Oil Corp	0.00053	Royal Dutch Petroleum	0.007632609	EUR
Unocal	0.0027923	TOTAL S.A.	0.006673918	US
Apogent Technology	0.0009996	Alcon Inc	0.000535088	EUR
Cleveland Cliffs	0.0010419	Rio Tinto	0.00936036	US
Carbo Ceramics	0.0043436	Core labs	0.002511999	EUR
Buckeye Tech	0.0008011	Stora Enso	0.000885683	EUR
Schweitzer Mauduit Intl	0.0020682	UPM-Kymmene Corporation	0.003924585	EUR
Bradley Pharmaceuticals	0.0020663	Aventis	0.005506266	US
Pharmaceutical Resources	0.0083422	GlaxoSmithKline plc	0.001110406	US
Medicis Pharmaceuticals	0.0035381	Novartis	0.002804125	EUR
Alpharma Inc	0.0009415	AstraZeneca Grp	0.002532697	US
Mylan Labs	0.0104231	Elan Corp	0.003581138	US
KV Pharmaceauticals	0.0043261	Schering Aktiengesellschaft	0.000825191	EUR
Allmerica Financial Corp	0.0023074	Wilis Grp.	0.004099447	US
Coachman Industries	0.008962	Pearson plc.	0.001165023	EUR
Dover Motorsports	0.0060444	Carnival Plc.	0.007122825	US
Memc Electronic Materials	0.0012676	Infineon Technologies	0.000452386	US
Fairchild Semiconductors	0.0029978	STMicroelectronics	0.000928274	US
Arch Chemicals	0.0015351	Imperial Chemical Inds. PLC	0.0005951	EUR
Rogers Corp	0.0018954	Syngenta	0.003973407	US
Cadence Design System	0.0018845	SAP	0.001285704	US
Standard Commercial Corp	0.0034847	Gallagher Group Plc	0.000819935	EUR

The first column of the table reports the list of U.S firms, the second column the NYSE ticker symbols and the third column the mean percentage bid-ask spread for the trading period. The columns four, five and six report the same for the European firms and the seventh column reports the variable "X". If X= "US" ("EUR") it means the U.S.(European) firm in that pair has a higher percentage bid-ask spread. The column header perspread denotes the percentage bid-ask spread.

Sample Period 2:30-3 PM

Examining the evidence for bid-ask spreads from afternoon data, the mean *perspread* for the U.S. sample is 0.003287 and 0.0029956 for the European sample. So the mean *perspread* for the European sample is smaller than that of the U.S. sample by 3 basis points. The associated P value is 0.4954. Out of each of the pairs of firms it was found that for 35 pairs of firms the European firm has a bigger *perspread* and for the rest of the 37 pairs the U.S. firm has a bigger *perspread.* So the evidence seems to suggest that the pattern of bid-ask spreads for the U.S. and the European firms becomes more homogeneous during the NYSE afternoon than in the morning. We found the correlation coefficient between the mean *perspread* and the number of trades to be -0.0928 for the U.S. sample and -0.0104 for the European sample The strength of the inverse relation between the bid-ask spread and liquidity that we obtained in the NYSE morning has also diminished during the afternoon.

Cross Sectional Regression on Perspread: 9:30-10 AM

We first examine the hypothesis that during the NYSE morning the European firms trade at a smaller bid-ask spread than the U.S. firms because of the presence of a substitute market. To test this, we ran a cross-sectional regression. The regression is specified as follows. The dependent variable is *perspread*, PS_i. It is regressed on the following dependent variables: Probability of informed trading PIN_i, a dummy d_i that takes the value 1 for a European stock and 0 for a U.S. stock, and the consolidated number of trades, $Trades_i$.The regression is done using White heteroskedasticity-consistent standard errors and covariance. Here i denotes firm. The estimation results are shown in the following regression equation:

$$PS_i = 0.002685 - 0.001976 * d_i + 0.010838 * PIN_i - 0.00000000317 * Trades_i + \varepsilon_1 \; where \; \varepsilon_i \sim N(\xi, \delta^2) \quad (2)$$
$$(0.0068) \quad (0.0003) \qquad (0.0002) \qquad (0.0067)$$

We find some interesting results in this regression. The significantly negative value, 0.001976 of the coefficient on the dummy d_i denotes that, after controlling for informed trading and liquidity, if we switch from a U.S. to a European firm, the value of *perspread* goes down by approximately 20 basis points. The significantly positive coefficient value of 0.010838 on the variable PIN_i suggests that as the value of the probability of informed trading in any stock goes up by 1, the percentage bid-ask spread for that stock goes up by 108 basis points after controlling for the dummy and the number of trades. The significantly negative coefficient on the liquidity measure $Trades_i$ denotes that after we control for informed trading and dummy, percentage bid-ask spread goes down by 7 basis points as we increase the standard deviation of $Trades_i$ by one unit.

Cross sectional regression on Perspread: 2:30-3 PM

The effect on percentage bid-ask spread due to the presence of a substitute market should disappear when the European markets close. The last European market to close trading for our sample is floor trading at the Frankfurt Borse which closes trading at 8 pm local time in Frankfurt. This translates to 2 pm local time in New York. So from 2-4 pm local time in New York only the NYSE is trading. To validate our theory of smaller bid-ask spread in the presence of a substitute market, empirically, we should expect to see the difference in bid-ask spreads between the U.S. and the European firms disappear during the NYSE afternoon when all the European markets have closed. So we estimated the same cross-sectional regression as before using data from the time period 2:30-3pm using White heteroskedasticity-consistent standard errors and covariance. The results of the estimation analysis are shown in the following regression equation:

$$PS_i = 0.003669 - 0.000313 * d_i - 0.000906 * PIN_i - 0.000000000988 * Trades_i + \varepsilon_1 \; where \; \varepsilon_i \sim N(\rho, \psi^2) \quad (3)$$
$$(0.000) \quad (0.4782) \qquad (0.5024) \qquad (0.1666)$$

The results indicate that the coefficient on the dummy is negative, still, but not significant, which suggests that the U.S. and the European firms do not have significantly different percentage bid-ask spreads after we control for liquidity and informed trading. The effect of liquidity and probability of informed trading on the perspread is no longer found to be significant. The results of the cross sectional regression for the two periods 9:30-10 am and 2:30-3 pm are documented in Table 5 shown below.

Table 5: Cross Sectional Regressions

Panel A: Results for 930-10:00 AM				
	Coefficient	**Standard Error**	**t-statistics**	**Probability**
Constant	0.002685	0.0068	3.502580	0.0006***
Dummy	(-)0.001976	0.0003	-3.830808	0.0002***
PIN	0.010838	0.0002	5.075426	0.0000***
Trades	(-)0.00000000317	0.0067	-2.642661	0.0092***
R square	0.324213			
Panel B: Results for 2:30-3:00 PM				
	Coefficient	**Standard Error**	**t-statistics**	**Probability**
Constant	0.003669	0.000	10.02772	0.0000***
Dummy	(-)0.000313	0.4782	0.453134	0.6512
PIN	(-)0.000906	0.5024	-1.125026	0.2625
Trades	(-)0.00000000098	0.166	-1.000537	0.3188
R square	0.009403			

The table summarizes the regression results. The results for the time period 9:30-10 am and 2:30-3pm are presented in Panels A and B respectively. The estimates are followed by the standard error, t statistic and p values. We are using 5% level of significance. The R square value is reported at the end of the table.

CONCLUSION

The study has tried to answer the question of whether the market maker at the NYSE faces more competition for European stocks which have a substitute market open during morning trading hours in New York than for U.S. stocks This effect of the presence of a substitute market should disappear when all the European markets close trading during the NYSE afternoon. The results from our cross-sectional regression analysis seem to support the hypothesis nicely as percentage bid-ask spreads of European stocks are smaller than that for U.S. stocks in the New York morning. This difference disappears during the New York afternoon when European trading has ended. This indicates that the U.S. and the European markets are integrated during the period of overlap.

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LABOR MARKET EFFICIENCY IN POLAND: A STOCHASTIC WAGE FRONTIER ANALYSIS

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ABSTRACT

In this paper, we apply a stochastic frontier approach in order to analyze labor market efficiency in Poland – a transition economy and a new entrant to the European Union. Wage efficiency is defined as the ratio of a worker's actual and potential (maximum) wage, given his or her demographic and socio-economic characteristics. Our findings indicate that, on average, in 2001 the full-time hired Polish workers realized 86 percent of their potential earnings. It implies inefficiency in acquiring and processing job market information. At the same time, an international comparison shows that the degree of wage efficiency in Poland was similar to or higher than that in other developed countries. Our attempt to identify the determinants of wage efficiency in Poland produced mixed results. However, in sum, worker performance in the Polish labor market seems to be rewarded appropriately, with some typical-for-Europe degrees of inefficiency in acquiring information, by a standard of wage efficiency and proximity to the wage frontier.

INTRODUCTION

The transition process to a market economy in the Central and East European countries has included a wide spectrum of adjustments in labor markets. Different aspects of these adjustments – such as unemployment, labor market flows, the wage structure and distribution, etc. – in post-Communist economies have been extensively scrutinized and analyzed in the economics literature. A significant body of research has focused on labor market developments in Poland, which is widely regarded as one of the most successful transition economies in Central and Eastern Europe. In particular, a number of empirical studies have analyzed the dynamics of wage distribution and wage structure in Poland during the pre-transition years, and then the early and mature stages of transition (see Adamchik *et al*., 2003 for an overview). A frequently used approach is the Mincerian earnings function (Mincer, 1974). However, this model assesses the impact of different factors on the *average* level of earnings; it does not identify the potential (maximum) earnings level for a worker with a given set of worker characteristics and how these characteristics contribute to achieving the potential wage.

In this study, we apply a stochastic frontier approach in order to analyze the degree of wage efficiency in the Polish labor market in 2001. To our knowledge, there has been no such research for Poland. Wage efficiency is defined as the ratio of a worker's actual and potential wage, given his or her demographic and socio-economic characteristics. Consequently, wage inefficiency is defined as the gap between a worker's actual and potential wage. Wage inefficiency arises mainly from incomplete labor market information and inefficient job matches, which results in the loss of output. Thus, understanding the causes and extent of wage inefficiency in a country will help to develop more appropriate labor market policies and institutions, ultimately to increase national output. We analyze the factors that affect the potential wage, and focus on the degree of wage (in)efficiency and its determinants. The paper is organized as follows: The Methodology section sketches the key concepts for a framework of stochastic frontier models. The Data section explains the data set that we use for our estimates of the frontier and resulting wage efficiency ratios. In the next two sections, we investigate factors that influence the wage frontier and analyze wage efficiencies at various levels of disaggregation. The International Comparison section provides a perspective from existing study with which to interpret our results for Poland. The final section concludes the paper.

METHODOLOGY

In a labor market, worker i with a given set of demographic and socio-economic characteristics faces a wage-offer distribution varying from the lowest wage ($w_{\min i}$) to the highest potential wage ($w_{\max i}$). Workers whose actual wage (w_i) is less than their potential maximum wage are said to be suffering from some kind of "wage inefficiency." Inefficiency may be attributed to different causes, such as imperfect information on the part of both workers and employers, discrimination, the market power of the employer, or a worker's negotiating power. For instance, while searching for a job, workers do not know which firms pay the highest wages for their set of skills. Because search is costly, workers may stop searching and accept lower wages before discovering the highest-paying job. On the other hand, employers have imperfect information about potential hires. Different employers have access to different data about the same worker, leading to different conclusions about this person and different wage offers to him or her.

The earnings frontier approach describes the maximum potential wage for a worker with a specific set of characteristics. All workers are assumed to be located either on this "best available wage," envelope frontier (a fully efficient position when $w_i = w_{\max i}$) or below this frontier (an inefficient position when $w_i < w_{\max i}$). Early deterministic frontier models (Greene, 1980) assumed that each deviation from the frontier (i.e., potential wage $w_{\max i}$) was due to inefficiency:

$$\ln w_i = \ln w_{\max i} - \varepsilon_i , \qquad (1)$$

where $\ln w_i$ and $\ln w_{\max i}$ are the logarithms of the observed and potential wage of the *i*-th individual; and ε_i is a one-sided non-negative error term, because of the impossibility of $w_i > w_{\max i}$. Aigner *et al*. (1977) proposed that the frontier itself may be stochastic and split the error term into two parts – a white noise variable $v_i \sim N(0,\sigma_v^2)$ and a non-negative inefficiency term $u_i \geq 0$. The wage frontier is usually modeled with a Mincerian earnings function (Mincer, 1974). Eqn. (1) thus can be rewritten as:

$$\ln w_i = \ln w_{\max i} + v_i - u_i = \alpha + \boldsymbol{\beta}'\mathbf{x}_i + v_i - u_i \ , \qquad (2)$$

where $\mathbf{x}_i$ is a vector of socio-economic characteristics; and α and β's are parameters to be estimated. In Eqn. (2), $\ln w_{\max i}$ represents a deterministic wage frontier, $\ln w_{\max i} + v_i$ represents a stochastic frontier, and $\ln w_{\max i} + v_i - u_i$ represents the observed wage. The degree of wage inefficiency for each worker is measured by the difference between the actual wage and the stochastic wage frontier (that is, u_i). In this study we assume that wage inefficiency generally results from the imperfect information of employees. However, as mentioned above, wage inefficiency may be attributed to incomplete information of both workers and employers. Polachek and Yoon (1987, 1996) proposed such a model with a two-tiered stochastic wage frontier, in which the error term ε_i is split into three parts: white noise, a non-positive error term representing worker information gaps, and a non-negative error term for employer ignorance. However, as Polachek and Xiang (2005, p.7) later recognized, empirical results suggest that "incomplete employee information varies far more than incomplete employer information" and thus may be ignored without a significant loss of accuracy and generality.

We assume a half-normal distribution for $u_i \sim N^+(0,\sigma_u^2)$ and that u_i and the independent variables are unrelated. Eqn. (2) is estimated using the log-likelihood function (Aigner *et al*., 1977; Meeusen and van den Broeck, 1977). The conditional expected value of u_i given ε_i is calculated as in Jondrow *et al.* (1982). Finally, we use u_i-values to calculate individual efficiency (EFF) and inefficiency (INEFF) ratios which measure the gap between the actual wage and the stochastic wage frontier:

$$\mathrm{INEFF_i} = 1 - \mathrm{EFF_i} = 1 - \frac{\exp(\alpha + \boldsymbol{\beta}'\mathbf{x}_i + v_i - u_i)}{\exp(\alpha + \boldsymbol{\beta}'\mathbf{x}_i + v_i)} = 1 - \exp(-u_i)\,. \qquad (3)$$

We then explore whether a set of macroeconomic, demographic, socio-economic, and institutional characteristics can explain the variation of the efficiency estimates.

Data

The Labor Force Survey conducted by the Polish Central Statistical Office in May of 2001 constituted the data source for this paper. We restricted our attention to full-time hired workers because only this category of employees was required to report their net earnings at their main workplace during the preceding month. Part-time hired workers, self-employed individuals, and those assisting in family businesses were not required to report their earnings. We further narrowed our sample by deleting those individuals who were full-time students, or handicapped, or younger than 18, or older than 65 (men) and 60 (women). These age restrictions correspond to the specific retirement ages as well as comply with the Polish Central Statistical Office definition of the working-age population (women 18-59 women; men 18-64). Furthermore, because wages were defined in terms of monthly earnings, for consistency we controlled for an employee who worked 40 and more hours per week on a regular basis. After all these adjustments, we had a sample of 9,380 full-time hired employees, of which 5,208 were males and 4,172 were females.

ESTIMATES OF THE STOCHASTIC WAGE FRONTIER

The maximum likelihood estimates of Eqn. (2) and the means of independent variables are presented in Table 1. In addition to the conventional human capital characteristics (education, potential experience, gender), our wage frontier equation includes other personal characteristics (marital status, head of the household) as well as dummy variables that capture regional labor market conditions (region, city/town size or village). We also include current job characteristics such as tenure, which reflects years of work experience with the current employer, as well as controls for thirteen industries, eight occupational indicators, four firm sizes, and an indicator for the sector of work (public versus private). Similar to other studies in this area, we use potential experience (age minus years of completed education minus 6) as a proxy for actual experience. For the regression model in Table 1, the reference person is a woman who has an elementary or lower education, is not married, does not head the household, lives in a small town or rural area in the Central region, works as a laborer in a small (5 or fewer employees) private manufacturing firm, and has less than a year of both potential experience and tenure at the current workplace.

The results in Table 1 show that the potential wages of Polish workers are closely related to their demographic and socio-economic characteristics. Most variables in Table 1 are significant at the 5-percent level or less and have the anticipated signs. For instance, the coefficients of the gender, education, and potential experience variables are statistically significant and positive and indicate that the potential wages were higher for men, and for workers with more education or potential experience. According to our results, men have a 15.2 percent higher potential wage than women, *ceteris paribus*. University-educated workers experience a 38.0 percent higher potential wage than their counterparts with only elementary education. The impact of potential experience on the wage frontier exhibits a standard concave shape: each additional year of potential experience increases the potential wage but at a decreasing rate, so that after about 35 years the positive impact of an additional year of experience starts to decline. The impact of tenure is similar to that of potential experience. Being married increases the wage potential by 4.4 percent; however, we should treat this result with caution. It is well documented in the literature that marriage positively affects men's wages, but inversely for women. Thus, the "married" coefficient in Table 1 may be misleading. Further, such factors as living in a more economically developed Central region (including Warsaw), or in a big city, or being a top manager, or working in a firm with more than 100 employees indicate a larger potential wage. Workers in the public sector face a lower potential wage than their similarly endowed counterparts in the private sector. These results are quite consistent with economic theory and with similar studies that have estimated the wage frontier for different countries.

Table 1: Maximum Likelihood Estimates of the Stochastic Wage Frontier

Variable	Coeff.	Std.Err.	Mean	Variable	Coeff.	Std.Err.	Mean
Constant	6.395*	(0.030)		Transportation	0.069*	(0.013)	0.081
Man	0.152*	(0.008)	0.555	Financial interm.	0.190*	(0.024)	0.022
University	0.380*	(0.019)	0.100	Real estate	-0.004	(0.016)	0.038
Post-secondary	0.162*	(0.020)	0.043	Public admin	0.054*	(0.015)	0.077
Secondary vocational	0.144*	(0.014)	0.295	Education	-0.100*	(0.020)	0.041
Secondary general	0.155*	(0.017)	0.075	Health care	-0.153*	(0.016)	0.089
Basic vocational	0.054*	(0.012)	0.384	Social services	-0.018	(0.020)	0.029
Potential exp., years	0.007*	(0.001)	20.750	Top manager	0.471*	(0.019)	0.040
Potential exp. sq.	-0.000*	(0.000)	538.313	Specialist	0.311*	(0.019)	0.066
Married	0.044*	(0.008)	0.747	Technician	0.270*	(0.015)	0.163
Head of household	0.104*	(0.007)	0.508	Office clerk	0.131*	(0.016)	0.106
Region South	-0.038*	(0.011)	0.147	Services	0.070*	(0.017)	0.117
Region East	-0.114*	(0.010)	0.214	Farmer	0.058	(0.044)	0.009
Region North-West	-0.038*	(0.010)	0.188	Industrial worker	0.089*	(0.013)	0.254
Region South-West	-0.035*	(0.012)	0.114	Machinist	0.126*	(0.014)	0.134
Region North	-0.065*	(0.010)	0.186	Public sector	-0.021*	(0.009)	0.429
City (>100 thous.)	0.069*	(0.020)	0.280	Tenure, years	0.008*	(0.001)	10.047
City (50-100 thous)	-0.003	(0.021)	0.101	Tenure sq.	-0.000*	(0.000)	189.053
City (20-50 thous.)	-0.025	(0.021)	0.125	Firm size (6-20)	0.052*	(0.013)	0.195
City (10-20 thous.)	-0.030	(0.021)	0.089	Firm size (21-50)	0.083*	(0.014)	0.156
City (5-10 thous.)	-0.053*	(0.024)	0.036	Firm size (51-100)	0.087*	(0.014)	0.139
Rural	-0.023	(0.020)	0.336	Firm size (>100)	0.129*	(0.013)	0.424
Agriculture	-0.047*	(0.020)	0.026				
Mining	0.226*	(0.023)	0.024	λ	0.701*	(0.047)	
Energy supply	0.096*	(0.022)	0.030	σ	0.332*	(0.003)	
Construction	0.059*	(0.012)	0.082	σ_v^2	0.074		
Trade	0.006	(0.011)	0.137	σ_u^2	0.036		
Hotel & restaurants	0.047	(0.025)	0.018	N obs.	9380		

** Significant at the 5 percent level or less.*

ESTIMATES OF WAGE EFFICIENCY

The degree of asymmetry of the disturbance term is measured by λ, defined in Eqn. (4),

$$\lambda = \sigma_u / \sigma_v \quad . \tag{4}$$

While the estimate of λ in Table 1 is statistically significant, its low magnitude of 0.701 indicates that the inefficiency component in the data is rather small. For comparison, some studies report much higher values of λ: 1.06 for Germany (Lang, 2004), 1.83 for the U.S. and 2.65 for Canada (McClure *et al*., 1998).

We further decompose the variance of the composite error ε_i and calculate the contribution of the variance of u_i to the total variance. As Greene (1993) points out, for the half-normal model

$$\frac{\mathrm{Var}(u_i)}{\mathrm{Var}(\varepsilon_i)} = \frac{[(\pi-2)/2]\sigma_u^2}{\sigma_v^2 + [(\pi-2)/2]\sigma_u^2} \quad . \tag{5}$$

Substituting the estimates for each statistic from Table 1 into Eqn. (5), we find that only about 22 percent of the variance of ε_i results from wage inefficiency, and the remaining 78 percent is due to other unexplained variability factors. This finding reinforces our conclusion above that wage inefficiency plays a quite limited role in our estimates. We also note that our result is consistent with the magnitude of the U.S. estimate of 27 percent reported by Hunt-McCool and Warren (1993). In contrast, Lang (2004) reported an estimate of 39 percent for Germany, and Landeau and Contreras (2003) 51 percent for Chile. Although larger, these latter two estimates are within the same order of magnitude as ours for Poland and may indicate a somewhat different operation of those national labor markets, but not a completely different structure.

We now turn to the interpretation of the efficiency ratios. The estimated efficiency ratios, based upon Eqn. (3), for the entire sample and for different socio-demographic groups are presented in Table 2. For the entire sample, the efficiency ratio is 86 percent, that is, on average workers realize 86 percent of their potential earnings and are 14 percent below their potential. It means that an average worker could increase his or her wage by about 16 percent (1/0.86-1=0.16) without any additional investment in his or her human capital endowment. The efficiency ratio of 86 percent for Poland is quite consistent with the reported results for some countries. For example, 86 percent for the U.S. (Hunt-McCool and Warren, 1993), 84 percent for the U.S. and 83 percent for Canada (McClure *et al*., 1998), 83 percent for Chile (Landeau and Contreras, 2003), 80 percent for Germany (Lang, 2004), 80 percent for the UK (Polachek and Xiang, 2005).

Our next step is to determine whether wage efficiency varies among socio-demographic groups. The common rationale is that higher costs of job search, weak labor market attachment, environment with limited public knowledge, etc. lead to less complete information and, consequently, to higher wage inefficiency. Thus, typical expectations are that men, married workers, prime age workers, workers with more education, workers in urban areas, and natives experience less underpayment as compared to women, single workers, young workers, less educated workers, workers in rural areas, and migrants. For instance, the greater market attachment of men is believed to result in their having better labor market information and higher wage efficiency as compared to women (Groot and Oosterbeek, 1994). Residing in rural areas increases information costs and is likely to result in higher wage inefficiency as compared to the areas with dense population. Because migrants in the labor market usually possess less information than the native population, the former are expected to experience higher wage inefficiency (Polachek and Xiang, 2005). Contrary to these expectations, we fail to detect significant differences in wage efficiency for the above-mentioned and other population groups (see Table 2). While surprising, such results are not unusual. For instance, Lang (2004) does not detect any difference in wage efficiency between natives and immigrants in Germany; Dar (2006) reports no difference in wage efficiency between Canadian men and women and only a slight advantage for the native-born and university-educated Canadians. Possibly, these results occur from only analyzing the marginal effects of single differences, but more complex comparisons of joint pairs of characteristics might yield statistically significant results. Also for Poland, the extreme upheavals of the transition decade may have sensitized all workers in a similar fashion to information about wages.

Table 2: Wage Efficiency (Percent of the Wage Frontier) by Socio-Demographic Groups

Group	Mean	Std. Dev.	Min.-Max.	N obs.	Group	Mean	Std. Dev.	Min.-Max.	N obs.
All sample	86.0	3.8	42.5-95.4	9380	Industrial worker	86.1	3.8	47.6-95.0	2381
Men	86.0	4.0	42.5-95.4	5208	Machinist	86.0	3.9	48.9-95.4	1259
Women	86.1	3.6	42.8-94.9	4172	Manual worker	86.2	3.4	48.7-94.0	1037
University	85.7	4.7	42.8-94.4	941	Public sector	86.1	3.7	42.5-94.4	4023
Post-secondary	86.0	3.7	73.5-93.8	403	Private sector	86.0	3.9	46.7-95.4	5357
Secondary vocat.	86.1	3.6	47.8-94.4	2764	City (>100 thous.)	85.9	4.2	42.8-95.0	2622
Secondary gen.	86.0	4.0	46.7-94.9	704	Town & rural	86.1	3.7	42.5-95.4	6758
Basic vocat.	86.1	3.7	42.5-95.0	3600	Potential exp. <=10yrs	86.0	3.9	46.7-94.6	2063
Elementary	86.1	3.9	48.9-95.4	968	Potential exp. >10 yrs	86.0	3.8	42.5-95.4	7317
Married	86.0	3.8	42.5-95.4	7011	Firm size (>100 empl.)	86.0	3.9	42.5-94.3	3978
Not married	86.0	3.9	46.7-94.6	2369	Firm size (<100 empl.)	86.0	3.8	46.7-95.4	5402
Top manager	85.7	5.2	42.5-93.6	377	Region South	86.1	3.6	68.6-94.0	1378
Specialist	85.7	4.9	42.8-94.4	620	Region East	86.1	3.5	48.9-95.4	2005
Technician	86.0	3.8	58.8-94.9	1532	Region North-West	86.1	3.8	42.5-94.4	1765
Office clerk	86.1	3.6	46.7-94.3	991	Region South-West	86.0	4.1	42.8-94.9	1071
Services	86.2	3.2	52.7-94.6	1100	Region North	86.0	4.0	46.7-95.0	1749
Farmer	86.2	3.1	76.5-93.5	83	Region Central	85.9	4.1	58.8-93.9	1412

We then looked for a set of selected macroeconomic, demographic, socio-economic, and institutional characteristics that would explain the variation of the efficiency measures. We considered the 16 Polish administrative regions (voivodships), each of which had distinctive labor market characteristics. For each of these regions, we collected specific macroeconomic indicators that we believed could affect an individual's incentives to search for a higher wage and influence the acquisition of additional labor market information. Following Polachek and Xiang (2005), we tested population density, rural population, industrial employment, public sector employment, and the inflow of workers (both from other Polish regions and from abroad). We regressed the logarithms of these variables on the logarithm of the average regional wage efficiency ratio (EFF) defined in Eqn. (3). The estimation results are presented in Table 3.

The "population density" and "public sector employment" coefficients are both positive, which is consistent with economic theory. In the former case, a more dense population implies better access to information as well as more concentrated job opportunities (Sandell, 1980). In the latter case, as Groot and Oosterbeek (1994, p. 388) contend, "workers in the public sector possess more market information than workers in the private sector (…) probably due to the fact that wage policies in the private sector are in general less public knowledge and more individually based." Also, as expected, the "rural population" coefficient is negative because "rural regions are less concentrated with job opportunities, and therefore likely to result in more incomplete information" (Polachek and Xiang, 2005, p. 17). Our estimation results confirm that the two factors - population density and public sector employment - prolong search by lowering its costs, which in turn leads to gathering more information and increasing wage efficiency; and residing in rural areas has exactly the opposite effect on wage efficiency. While the signs of these three coefficients are in accord with job search theory, their statistical significance is weak, possibly due to a relatively high pairwise multicollinearity among industrial, public and rural employment variables. On the other hand, the "industrial employment" and "inflow of workers" coefficients are statistically significant, but their signs are opposite those expected from other countries' explanations of efficiency relationships. According to Freeman (1980), Polachek and Yoon (1987), Polachek (2004), and Polachek and Xiang (2005), the expected sign is positive for the "industrial employment" coefficient and negative for the "inflow of workers" coefficient. Industrial workers are assumed to be more strongly unionized, leading to more current information on wages and jobs, compressed wage distributions, and an increase in the degree of workers' wage efficiency; and migrants (both internal and from abroad) are assumed to have less knowledge than natives about the distribution of wages in the new region, with consequent higher wage inefficiency. However, this reasoning may not apply to our case because Poland has experienced a dramatic drop in unionization over the 1990s, along with major redistribution of former industrial workers, leaving "survivors" in industry less able to bargain for higher wages. Furthermore, a large portion of migrants in Poland is internal and may be highly sensitized to the opportunities opened by transition processes – they are simply following their new incentives with better labor mobility. Overall, using voivodship characteristics as the basis to explain regional wage efficiencies provides some empirically sensible (although statistically weak) results.

Table 3: Impact of Regional Macroeconomic Variables on the Average Regional Wage Efficiency Ratio [a]

Variable	Coef.	Std.Err.	Mean**
Constant	4.490	(0.096)	
Ln (Population density, persons per sq. km)	0.005	(0.004)	129.6
Ln (Rural population, % of total population)	-0.002	(0.010)	40.1
Ln (Industrial employment, % of total employment)	-0.021*	(0.011)	24.9
Ln (Public sector employment, % of total employment)	0.006	(0.021)	25.2
Ln (Inflow of workers, % of total employment)	0.013*	(0.007)	0.7
R-squared, %	37		
N obs.	16		

** Significant at the 7-8 percent level.*
*** Means of the original variables, not their logarithms.*
[a] The estimation method is OLS. The dependent variable is LnEFF; the EFF ratio is defined in Eqn. (3).

INTERNATIONAL COMPARISONS

In this section we provide a more detailed cross-country analysis of wage efficiency. We assess how Poland fits into a group of eleven countries (ten OECD countries and Israel) for which a recent study by Polachek and Xiang (2005) is available. We estimated a wage frontier specification similar to that of Polachek and Xiang, which has a much smaller set of independent variables: years of education, potential experience, potential experience squared, and a dummy for gender (woman). The definitions of these variables in our study are identical to those in Polachek and Xiang's paper, and we can make some qualitative evaluations. Our estimates for Poland appear below theirs in Table 4.

As shown at the bottom of Table 4, the mean number of years of schooling for our Polish sample is 11.985, which is very similar to the means reported for most OECD countries (with the maximum of

13.288 for Canada, and the minimum of 9.492 for Ireland). The average potential experience for Poland is 20.750 years, which again is well within the OECD range – between the maximum of 29.905 (Ireland) and the minimum of 18.450 (Canada). On average, the proportion of women in our sample is 44.5 percent, which is only slightly lower than the 46-52 percent range for OECD countries. The estimated frontier coefficients for Poland are all significant at quite robust *p*-value levels and in strong agreement with those reported in Polachek and Xiang (2005). *Ceteris paribus*, one additional year of education increases the wage frontier by 8.5 percent in Poland and by 6.5-16.1 percent in OECD countries. Polish women face a 23.1 percent lower potential wage than men do with the same characteristics. In Polachek and Xiang's sample, the Netherlands has the lowest female disadvantage (5.8 percent) and Israel has the largest one (51.4 percent). Finally, potential experience exhibits a common concave shape: the positive impact of each additional year of experience on the wage frontier is initially increasing (but at a decreasing rate) and then decreases. One additional year of experience shifts the wage frontier up by 1.8 percent in Poland and by 1.8-4.3 percent in OECD countries. While the estimated wage frontier for Poland is very similar to those reported for OECD countries, the average efficiency ratios differ quite a bit. For Poland, the EFF ratio is 89.1 percent (this estimate is for the parsimonious specification of the wage frontier in Table 4; for our extended specification in Table 1, the EFF ratio is 86.0 percent). For OECD countries the EFF ratio ranges from 43.7 percent (Finland) to 79.6 (UK). Consequently, the INEFF ratio is 10.9 percent for Poland, lower than those for OECD countries.

Table 4: Maximum Likelihood Estimates of the Wage Frontier Coefficients: International Comparison, Polachek and Xiang (2005) Specification [a]

Country	Year	Years of schooling	Potential experience, years	Potential experience squared	Woman	Wage efficiency, %
Canada	2000	0.098	-0.006	0.0002	-0.314	66.0
Czech R.	1996	0.093	0.020	-0.0003	-0.321	72.6
Finland	2000	0.065	0.043	-0.0005	-0.421	43.7
Germany	2000	0.105	0.020	-0.0002	-0.187	64.3
Ireland	1996	0.091	0.038	-0.0004	-0.110	64.7
Israel	1997	0.128	0.042	-0.0005	-0.514	65.9
Netherlands	1999	0.065	0.026	-0.0003	-0.058	70.3
Norway	2000	0.073	0.046	-0.0008	-0.460	51.6
Sweden	2000	0.092	0.036	-0.0005	-0.392	52.9
UK	1995	0.161	0.018	-0.0001	-0.368	79.6
US	2000	0.116	0.027	-0.0003	-0.307	61.6
Poland:	2001					89.1
coefficient		0.085*	0.018*	-0.0002*	-0.231*	
std. error		(0.002)	(0.001)	(0.0000)	(0.007)	
means		11.985	20.750	538.313	0.445	
λ = 0.464* (std.err. 0.080); σ = 0.345* (std.err. 0.005); σ_v^2 = 0.098; σ_u^2 = 0.021; N obs.= 9380						

** Significant at a less than 1 percent level.*
[a] Authors' computations for Poland. For other countries - Polachek and Xiang (2005), Table 2. Polachek and Xiang estimated wage frontier equations for 10 OECD countries and Israel over a number of years. We are using the most recent year for each country from their study.

Our last step was to repeat the graphical analysis of Polachek and Xiang (2005), not presented here, in order to see where our 2001 Polish results would fit in their scatter diagrams thereby indicating how close Poland was to the typical measure of wage inefficiency in OECD countries. We combined our estimate of wage inefficiency (about 11 percent) with Polish Central Statistical Office measures of population density, rural population, industrial employment, and inflow of foreign workers (both in absolute and relative values). The resulting pairs of numbers were plotted within Polachek and Xiang's Figures 2 to 6. Given that the estimated wage inefficiency ratio for Poland is lower than those for OECD countries, it was not a surprise that in all five cases Poland appears to be an outlier, quite similar to the UK and the Czech Republic that exhibit the lowest inefficiency ratios of Polachek and Xiang's estimates (2005, pp.

24, 29-31). This leads us to speculate that these countries have some unique characteristics that could be the subject of productive future research with regard to wage efficiency. At the same time, we should treat Polachek and Xiang's low wage efficiency results with caution, because they are rather at the low end of the reported wage efficiencies. Many other studies find much higher wage efficiencies – about 80-85 percent, which are more consistent with our findings.

CONCLUSIONS

In this paper, we have applied a stochastic frontier approach in order to analyze the degree of wage efficiency in the Polish labor market in 2001, that is, after more than a decade of transition adjustments and three years prior to joining the European Union. Our findings indicate that full-time hired Polish workers realized 86 percent, on average, of their potential earnings. An international comparison shows that the degree of wage efficiency in Poland is high and quite similar to other developed countries. Our attempt to identify the determinants of wage efficiency in Poland produced mixed results for our specific choice of explanatory variables. However, in sum, the transformed labor market structure in Poland appears to value a sensible relationship between worker skills or attributes and wages paid, similar to other developed economies. Worker performance seems to be rewarded appropriately, with some typical-for-Europe degrees of inefficiency in acquiring information, by a standard of wage efficiency and proximity to the wage frontier.

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PRICE REACTION TO DIVIDEND INITIATIONS AND OMISSIONS IN EMERGING MARKET: EVIDENCE FROM PRE AND POST MARKET CRISIS IN BANGLADESH

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ABSTRACT

Dividend signalling and information content of dividends are areas of interest in financial literature. A vast majority of the research conducted on information content of dividend. However, no study has examined the effectiveness of dividend announcements as a signalling device in the stock market of Bangladesh. This study employs conventional event study methodology to investigate whether dividend announcements convey information to the market or whether investors dividend announcements as the signalling device of the firm's prospects. The analysis is completed for the time period before and after the1998 market crisis in Bangladesh. The sample consists of cash dividend announcements for Dhaka Stock Exchange (DSE) listed firms preceding and following the market crisis. The empirical results suggest that the reactions to dividend announcements are not significant either preceding or following the financial crisis in Bangladesh, therefore, announcements of dividends neither convey information to the market nor do investors consider dividend announcements as a signal.

INTRODUCTION

Numerous studies conducted in different countries have documented that the announcement of changes in dividends and earnings convey specific information to the market (Pettit, 1972; Charest, 1978; Aharony and Swary, 1980; Woolridge, 1982 and 1983; Asquith and Mullins, 1983; Brickley, 1983; Divecha and Morse, 1983; Benesh *et al.* 1984; Dielman and Oppenheimer, 1984; Eades *et al.* 1985; Wansley and Lane, 1987; Aharony *et al.* 1988; Born, 1988; Ghosh and Woolridge, 1988; Healey and Palepu, 1988; Ghosh and Woolridge, 1991; John and Lang, 1991; Marsh, 1993; and Abeyratna *et al.* 1996). However, recent studies that have examined the simultaneous announcements by firms have discovered that the signal of dividends and earnings may either corroborate or contradict each other or, in consequence, influence the level of any abnormal returns, which are earned by investors (Kane *et al.* 1984; Easton, 1991; Eddy and Seifert, 1992). Nevertheless, previous empirical studies suggest that positive (negative) dividend change announcements produce positive (negative) common stock price changes (Asquith and Mullions, 1983; Healey and Palepu, 1988; and Michaely *et al.* 1995).

The price reaction to the announcements of dividends in the Dhaka Stock market of Bangladesh is likely to be different from developed markets. Therefore, this study attempts to investigate whether dividend announcement convey information to the security market of Bangladesh or whether investors in Bangladesh consider dividend announcements as the signaling device of firm's future prospects. After the financial crisis in Bangladesh market in 1998, there were a significant changes in institutional setting such as the introduction of online trading system and as well as Central Depository System (CDS) but there was no significant change in the legal framework as the controlling mechanism for the stock market. To compare the price reaction to dividend announcements in the preceding and following financial crisis and to test whether financial reform in the stock market of Bangladesh in 1998 brings any change in the market scenario, this study captures dividend initiations, omissions, and dividend maintaining announcements in the pre and post financial crisis in Bangladesh. The empirical results suggest that security prices do not react to dividend initiations, omissions or unchanged dividend announcements and financial reform does not help to improve the market scenario.

The rest of this paper is divided into four sections. The reviews of all the major theoretical and empirical evidence along with the critical evaluation for identifying the security price reactions to the announcements of dividends are included in section II. Section III contains the description of data and methodology of the empirical analysis. The empirical results are reported in section IV. The summary and the concluding remarks are incorporated in section V.

THEORETICAL BACKGROUND

Miller and Modigliani (M-M) (1961) provide the most comprehensive argument in support of the irrelevance of dividends. M-M maintained that dividend policy has no effect on the share prices of the firm, i.e., whether profit is paid as dividend or retained does not make any difference. Under the condition of perfect capital markets, rational investors, and the absence of tax discrimination, (i.e. between dividend income and capital gains), given the firm's investment policy, its dividend policy may have no influence on the market price of shares (Miller and Modigliani, 1966).

On the other hand, the bird-in-the-hand theory claims that stockholders prefer dividend payments to earnings; therefore, dividend policy is relevant to the value of shares. The leading proponents of the bird-in-the-hand theory (Gordon, 1962; and Lintner, 1962) view that stockholders value a dollar received in dividends more highly than a dollar of earnings retained. Gordon (1963) and Walter (1963) also support the dividend relevance doctrine.

Michaely et al. (1995) investigate both the immediate reaction to the initiation or omission of dividends and the long term post announcement price performance and their findings are quite consistent with prior empirical evidence (e.g., Asquith and Mullins, 1983; and Healey and Palepu, 1988) that dividend omission leads to price drops and prices increase as a result of dividend initiation.

Kalay and Loewenstein (1985) find that during a three-day period surrounding dividend announcement, the actual returns, on average, significantly exceed both the returns predicted by the market model and the average daily returns realized over a recent period. Nevertheless, they mention that the market reaction to dividend announcements is sluggish, i.e., the excess returns persist for up to four trading days after the announcement date. In a subsequent study, Eades *et al.* (1985), find that for the sub-sample of dividend announcements that are separated sufficiently from ex-dividend dates, there is no evidence of sluggishness. They confirm that the market reaction to dividend announcements is biased.

Bajaj and Vijh (1995) that the average excess returns to all dividend announcements increases as the firm size and stock price decreases presented different results. Their findings on the firm size and stock price effects suggest that the observed price reactions may be due to microstructure-based reasons. Market microstructure can affect stock prices during dividend announcement periods for two reasons: the spillover of tax-related trading around ex-dividend days and trading behavior related to the dissemination of dividend information. The summary of the major empirical studies on the security price reaction to dividend announcements are presented in Table 1.

Despite a vast majority of studies published on price reaction to dividend announcements in the developed markets, very few are in the emerging markets. However, most of those studies employed event study methodology but researchers applied a variety of approaches and considered different event study periods to analyze the data. Overall, the empirical results suggest that that positive dividend change announcements produce positive stock prices and vice versa.

Table 1: Major Studies on Price Reaction to the Announcements of Dividend

Author(s)	Data Set	Method Used	Findings Regarding Security Price Reaction
1. Aharony and Swary, 1980	384 dividend increasing, 47 dividend decreasing, and 2968 dividend maintained announcements for NYSE listed 149 industrial firms for the period of 1/1/1963 – 31/12/1976.	(1)Measurement of Abnormal Performance, and (2) Cumulative Effects of Abnormal Returns Approach of the event study methodology for the period of ±10 days.	1) Dividend increasing announcements: stock price increases. 2) Dividend decreasing announcement: stock price decreases. 3) Dividend maintained announcements: no change in stock prices.
2. Asquith and Mullins, 1983	All dividend initiation announcements of 168 NYSE listed firms for the period of 1954-1963.	(1) T-Test Approach of average excess return, and (2) Regression Approach of the event study methodology for the period of ±10 days.	Dividend initiation announcements: stock price increases and in general increases shareholders wealth.
3. Woolridge, 1983	317 dividend-increasing announcement and 50 dividend decreasing announcements of NYSE listed 225 firms for the period of 1970-1977.	Comparison Period Return Approach of the event study methodology for the period of ±10 days.	1) Dividend increasing announcements: stock price increases. 2) Dividend decreasing announcement: stock price decreases.
4. Fehr's et al. 1988	1015 dividend increasing, and 65 dividend decreasing announcements of US firms for the period of 1/1/1980 – 31/12/1984.	(1)Measurement of Abnormal Performance, and (2) Cumulative Effects of Abnormal Returns Approach of the event study methodology for the period of ±5 days.	1) Dividend increasing announcements: stock price increases. 2) Dividend decreasing announcement: stock price decreases.
5. Woolridge and Ghosh, 1988	408 announcements of dividend cut of NYSE listed 12 firms for the period of 1971-1982.	Comparison Period Return Approach of the event study methodology (period of ±1 Quarter.	Dividend cuts announcement: stock price falls.
6. Eddy and Seifert, 1992	Contemporaneous and non-contemporaneous dividend announcements of 1111 US firm for the period of 1983-1985.	(1) Mean Adjusted Return Approach, and (2) Regression Approach of the event study methodology for the period of -3 days and +1 day.	1) Price reaction to the joint announcement is significantly greater than just one single announcement. 2) Price reaction to the announcement of joint announcement is approximately twice that to a non-contemporaneous announcement. 3) Price reaction to joint contradictory announcement is not significant.
7. Dhillon and Johnson, 1994	61 dividend increasing, and 70 dividend decreasing announcements of NYSE listed firms for the period of 1/1/1978 – 31/12/1987.	Mean Adjusted Return Approach of the event study methodology for the period of ±10 days.	1) Dividend increasing announcements: stock price increases. 2) Dividend decreasing announcement: stock price decreases.
8. Michaely et al. 1995	561 cash dividend initiations and 887 cash dividend omissions announcement of NYSE listed firms for the period of 1964-1988.	Buy-and-hold strategy of the event study methodology for the period of ±1 day.	1) Dividend initiation announcements: stock price increases. 2) Dividend omission announcement: short-term price impact is negative.
9. Abeyratna et al. 1996	Dividend increase, decrease, and maintained announcements of 617 UK firms for the period of 1/1/1991 – 30/6/1991.	Measurement of Abnormal Performance (T-Test) Approach of the event study methodology for the period of ±1 day.	1) Dividend increasing announcements: stock price increases. 2) Dividend decreasing announcement: stock price decreases.
10. Impson, 1997	660-dividend decrease announcement of US unregulated firms (1974 – 1993) and 65 dividend decrease announcements of US public utility period of 1974 – 1993.	Regression Approach of the event study methodology for the period of ±1 day.	Dividend decrease by public utilities prompt stronger negative market reactions than similar announcements by unregulated firms.

The empirical part of this paper investigates the security price reaction to the announcement of dividends in an emerging market. The dividend announcements are divided into three categories: good news/dividend initiations, bad news/dividend omissions and no news/dividend maintaining announcements. An event study methodology is used considering four event periods (60, 30, 20, and 10 days preceding and following the announcement of dividends) to compare the mean abnormal returns between the observed period (preceding the announcement) and the comparison period (following the announcement) and to examine whether the abnormal returns preceding and following the announcements are significantly different from zero.

DATA AND METHODOLOGY

This section of the paper employs a conventional event study methodology to examine the stock price reaction to the announcement of dividends. The announcement day is defined as the event day (Day = 0), which is the day before the day on which dividend announcement news is published in the daily newspapers or in the daily stock price quotations. The observation periods are -60 days, -30 days, -20 days and -10 days of the event day; +60 days, +30 days, +20 days and +10 days of the event day are the comparison periods for the study.

Primarily, all of the listed companies of Dhaka Stock Exchange are considered as the population of this study for the period of 1988-2003. However, as because of financial crisis in Asian financial markets in 1997/98 and a great deal of speculation, Dhaka stock market crashed in 1998. An automated trading system replaced the traditional outcry trading and government reformed Security Exchange Commission (SEC) regulations to protect general investors and to ensure transparency in the securities market of Bangladesh, therefore, this study focused on the preceding (1988-1997) and following (1999-2003) market reform of Bangladesh. A part of the market data was collected from the Dhaka Stock Exchange price quotations, published and unpublished records of the Dhaka Stock Exchange, and the data channel (DataStream), and the rest of the data was collected from Dhaka Stock exchange database. The announcement dates are obtained from the Dhaka Stock Exchange daily price quotations for this study.

Daily share price returns are estimated according to the following equation (dividends are not included to estimate the stock returns):

$$R_{it} = (P_{it} - P_{it-1}) / P_{it-1} \tag{1}$$

Where,

R_{it} = Stock return on day 't'
P_{it} = Stock price on day 't' and
P_{it-1} = Stock price on day 't-1'

Abnormal returns are calculated according to the following equation:

$$AR_{it} = R_{it} - E(R_{it}) \tag{2}$$

Where,

AR_{it} = Abnormal return on day 't' and
$E(R_{it})$ = Expected return on day 't'

The expected return is derived using the well-known market model and based on the previous 300 days of the event study period.

Therefore, the expected returns '$E(R_{it})$' are calculated as:

$$E\left(R_{it}\right) = \hat{\alpha} + \hat{\beta} R_{mt} \tag{3}$$

Where,

$\hat{\alpha}$ = Predicted Value of Constant term
$\hat{\beta}$ = Predicted Value of Beta Coefficient, and
R_{mt} = Market return on day 't' {(Price Index_t – Price Index_{t-1}) / Price Index_{t-1}}

The Dhaka Stock Exchange index comprises both frequently and infrequently traded shares. However, it is also known that frequently traded shares cause upward bias and infrequently traded shares cause downward bias. Scholes and Williams (1977) and Dimson (1979) explained the problem of infrequent trading bias in the financial markets and mentioned the problem of using OLS model. They suggest considering lag and lead factor for adjusting upward and downward bias. On the other hand, Bartholdy and Allan (1994) considered Scholes and Williams (1977) and Dimson's (1979) suggested lag and lead factors alongside the OLS model but they found more stability of the coefficients in case of using the OLS model. Therefore, using the market model for predicting constant terms ($\hat{\alpha}$) and beta coefficients ($\hat{\beta}$) is quite justified for this study.

All cash dividend announcements of the listed firms of the Dhaka Stock Exchange over the period of 1988-2003 are primarily considered as the sample of the study. There were 801 cash dividend announcements in the sample period but 59 of them we excluded as the announcements for year 1998. Out of remaining 742 announcements, 232 cash dividend announcements are excluded because those announcements accompanied earnings and/or rights and/or stock dividend announcements and/or the announcements were made in the event study period. Therefore, the final sample consists of 510 cash dividend announcements amongst 352 announcements in the preceding and 158 in the following financial crisis in the stock market of Bangladesh. There are 198 dividend increasing announcements (initiations), 79 dividend-decreasing announcements (omissions), and 75 dividend maintaining announcements in the pre-crisis sample (1988-97) and 70 increasing (initiations), 46 decreasing (omissions), and 42 dividend maintaining announcements in the post-crisis sample (1999-03).

Hypothesis of the study:
H_0: *The mean abnormal returns of the observation period and comparison period are not significantly different from zero.*

The empirical part of this paper investigates the security price reaction to the announcement of increasing dividends (initiations), decreasing dividends (omissions) and maintaining dividends. To investigate the security price reaction to the announcement of dividends, the empirical part compares the abnormal returns of the observation and comparison period for four event study periods (±60 days, ±30 days, ±20 days and ±10 days) simultaneously.

EMPIRICAL EVIDENCE

We discussed the empirical results in this section. The discussion is broken down into three parts. The first part discusses the price reaction to good news or dividend initiations. The second part discusses the price reaction to bad news or dividend omissions. Finally, the third part discusses the price reaction to no news or dividend maintaining announcements.

Good News/Dividend Initiations

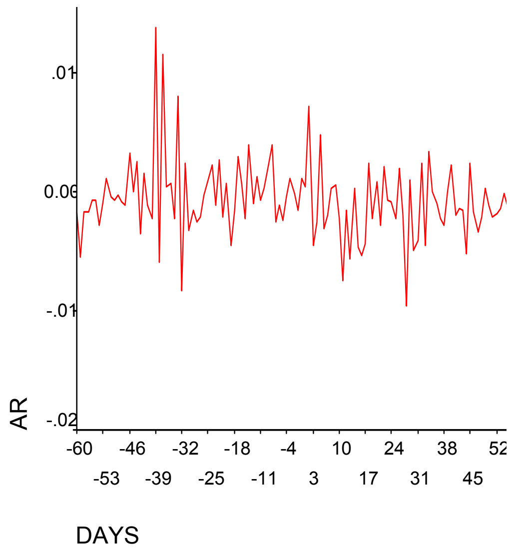
The mean abnormal returns in the pre-crisis sample are -.0087%, -.028%, -.0062% and -.010% in the observation periods -60 days, -30 days, -20 days, and -10 days respectively but these returns decrease in the comparison periods +60 days, +30 days, and +20 days, and +10 days (-.12%, -.16%, -.15%, and -.012). However, the mean abnormal returns in the post-crisis sample are -.20%, -.22%, -.28%, and -.31% in the observation periods -60 days, -30 days, -20 days, and -10 days respectively but these returns slightly increase in the comparison periods +20 days, and +10 days (-.24%, and -.28%) and decreases in +30 days (-.23%), but remains unchanged in +60 days(-.20%). Despite a slight decrease of returns after the increasing announcements in the pre-crisis sample, the effect of the announcement is mixed in the post-crisis sample; therefore, the signal of this sort of announcement is unclear (see table 2).

The correlation coefficients between abnormal returns of observation periods and comparison periods of the pre-crisis sample are -.232, -.076, -.243, and .116 and the probability values are .075, .690, .301, and .750 respectively for ±60 days, ±30 days, ±20 days and ±10 days. However, the correlation coefficients between abnormal returns of observation periods and comparison periods of the post-crisis sample are -.019, -.062, -.228, and -.108 and the probability values are .886, .745, .334, and .766 respectively for ±60 days, ±30 days, ±20 days and ±10 days. The correlation coefficients indicate a negative relationship between the abnormal returns of the observation periods and comparison periods for dividend initiations in all the study periods (±60 days, ±30 days, ±20 days and ±10 days) except pre-crisis ±10 days. Nevertheless, these results do not explain a high degree significant correlation between the abnormal returns of observation periods and comparison periods even in a single pair (see table 3).

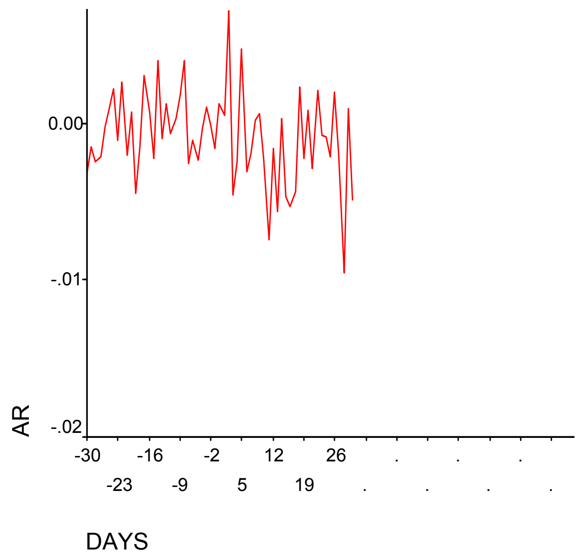
The mean difference between the abnormal returns of the observation and the comparison periods of the pre-crisis samples are .0012, .0013, .0014, and .0001 respectively for ±60 days, ±30 days, ±20 days and ±10 days. However, the mean difference between the abnormal returns of the observation and the comparison periods of the post-crisis sample are .0001, .0001, -.0004, and -.0003 respectively for ±60 days, ±30 days, ±20 days, and ±10 days. The t-values of the pre-crisis sample are 1.772, 1.681, 1.340, and .016 respectively. There probability values are .082, .103, .196, and .987 respectively for ±60 days, ±30 days, ±20 days and ±10 days. However, the t-values and the probability values of post-crisis sample are .023, .086, -.208 and -.105, and .981, .932, .837 and .919 respectively for ±60 days, ±30 days, ±20 days and ±10 days. These results failed to imply that the mean difference of the abnormal returns between observation and comparison periods is not significantly different from zero either in the preceding or following financial crisis sample (see table 4). Nevertheless, the sequence charts of the abnormal returns for the event study periods of ±60, ±30, ±20, and ±10 days (Figure 1 and 2) support the same argument. Therefore, the empirical evidence contradicts with the previous studies of price reactions to dividend initiations (see Table 4).

Figure 1: Good News/Dividend Initiations: Pre-crisis Sample (1988-1997)

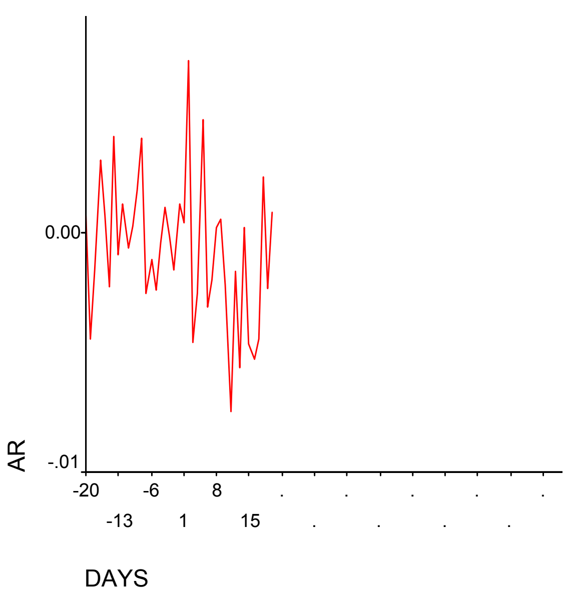
± 60 Days of the Event Day



± 30 Days of the Event Day



± 20 Days of the Event Day



± 10 Days of the Event Day

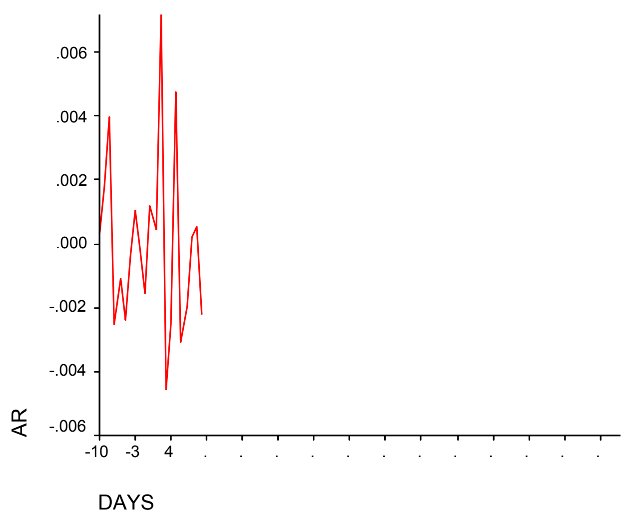
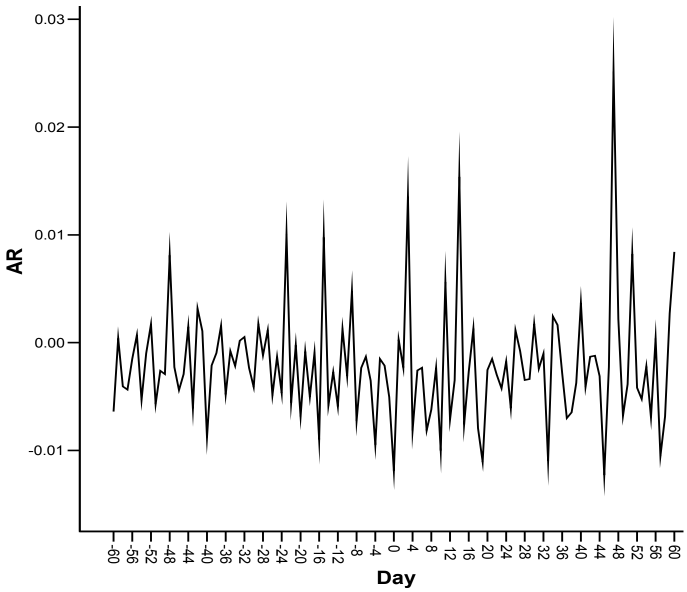
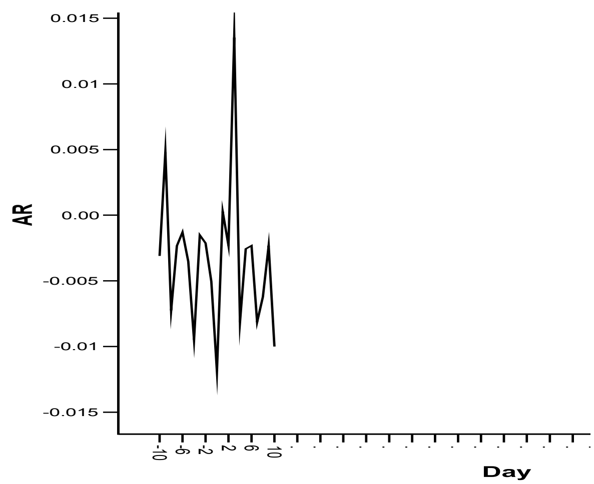


Figure 2: Good News: Post-crisis Sample (1999-2003)

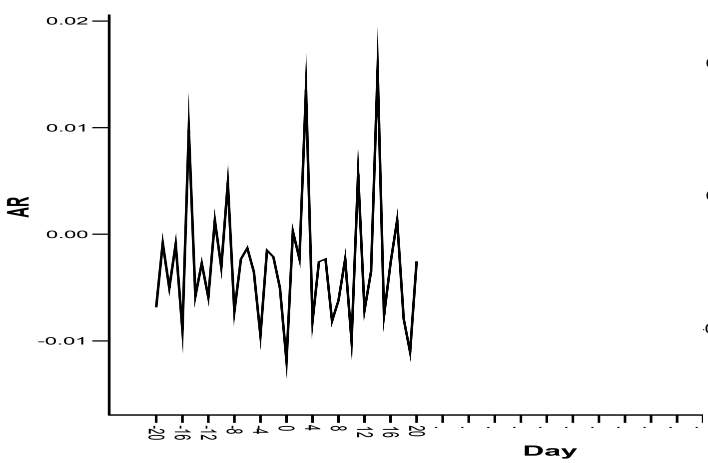
± 60 Days of the Event Day



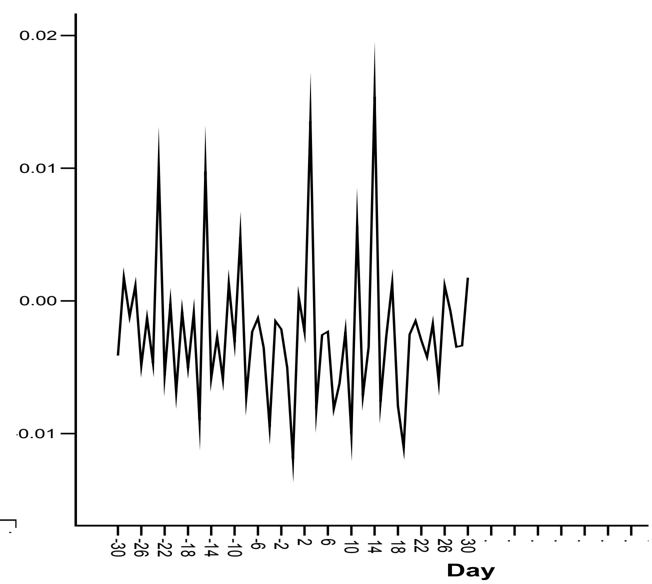
± 10 Days of the Event Day



± 20 Days of the Event Day



±30 Days of the Event Day



<u>Bad News/Dividend Omissions</u>

The mean abnormal returns in the pre-crisis sample are .75%, -.76%, -.77% and -.81% in the observation periods -60 days, -30 days, -20 days, and -10 days respectively, but these returns decrease in the comparison periods +60 days, +30 days, and +20 days, and +10 days (-.91%, -.89%, -.96%, and -1.03%). However, the mean abnormal returns in the post-crisis sample are .02%, .01%, -.03% and -.18% in the observation periods -60 days, -30 days, -20 days, and -10 days respectively, but these returns also decrease in the comparison periods +60 days, +30 days, +20 days, and +10 days (-.04%, -.31%, -.41% and -.76%). Despite a slight decrease in the returns after the bad news, the significance is irrelevant in terms of size after the decreasing dividend announcement (see table 2).

The correlation coefficients between abnormal returns of observation periods and comparison periods of the pre-crisis sample are .029, -.330, -.603, and -.630 and their probability values are .829, .075, .005, and .051 respectively for ±60 days, ±30 days, ±20 days and ±10 days. However, the correlation coefficients between abnormal returns of observation periods and comparison periods in the post-crisis sample are -.200, -.018, -.134, and -.222 and their probability values are .126, .926, .573, and .537 respectively for ±60 days, ±30 days, ±20 days and ±10 days. The correlation coefficients indicate a negative relationship between the abnormal returns of the observation periods and the comparison periods for the dividend decreasing announcements in all the study periods (±60 days, ±30 days, ±20 days and ±10 days) but failed to explain a very high level of significance in either pair (see table 3).

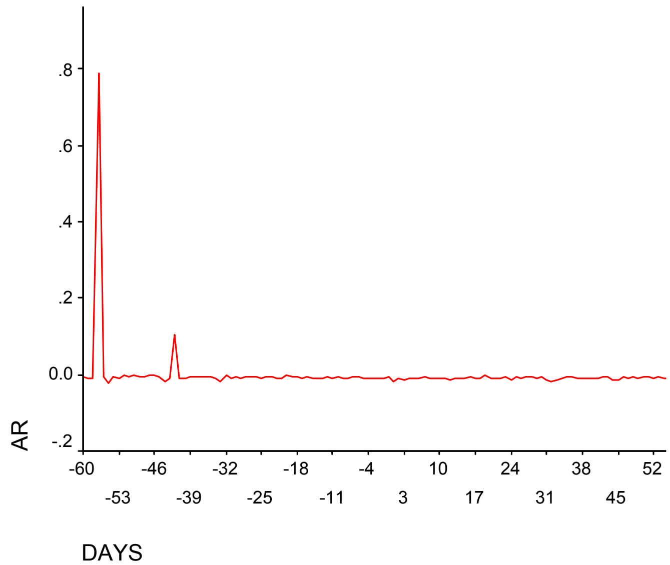
The mean difference between the abnormal returns of the observation and the comparison periods in the pre-crisis sample are .0166, .0013, .0019, and .0022 respectively for ±60 days, ±30 days, ±20 days and ±10 days. However, the mean differences between the abnormal returns of the observation and the comparison periods in the post-crisis sample are .0006, .0032, .0038, and .0058 respectively for ±60 days, ±30 days, ±20 days and ±10 days. The t-values and the probability values of the pre-crisis sample are 1.243, 1.559, 1.683, and 1.261, and .219, .130, .109, and .239 respectively for ±60 days, ±30 days, ±20 days and ±10 days. However, the t-values and the probability values of the post-crisis sample are .410, 1.577, 1.230, and 1.084 and their probability values are .683, .126, .234, and .307 respectively for ±60 days, ±30 days, ±20 days and ±10 days. These results, however, imply that the mean difference of the returns is not significantly different from zero. The sequence charts of the abnormal returns for the event study periods (Figure 3 and 4) also support the same argument. Nevertheless, the abnormal returns of the differential periods are not significantly different from zero. Despite the empirical results narrowly support the previous studies that dividend omissions produce negative stock prices (Asquith and Mullins, 1983; Healey and Palepu, 1988; and Michaely et al. 1995), the t-values are not significant in either pair in the current study, which makes the situation so ambiguous and indeed tough to come to a conclusion that security prices react negatively to dividend omissions (see table 4).

No News/Dividend Maintaining Announcements

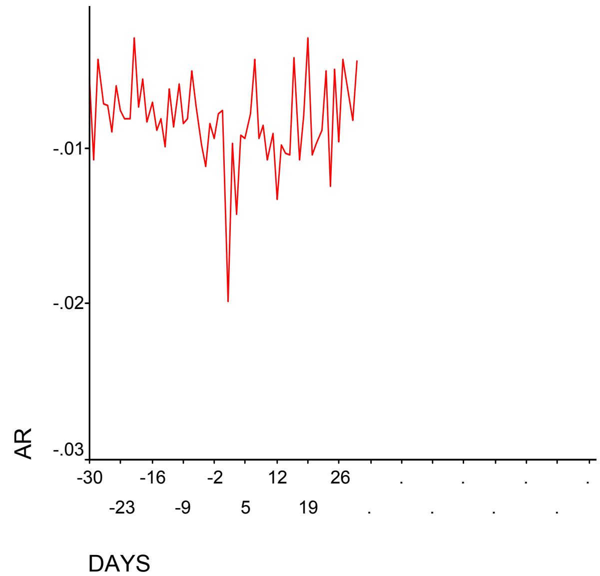
The mean abnormal returns in the pre-crisis sample are .04%, .05%, .06%, and .10% in the observation periods -60 days, -30 days, -20 days, and -10 days respectively, but these returns decrease in the comparison period +60 days, +30 days, and +20 days, and +10 days (-.051%, -.16%, -.20% and -.21%). However, the mean abnormal returns in the post-crisis sample are -.55%, -1.15%, -1.22%, and -1.16% in the observation periods -60 days, -30 days, -20 days, and -10 days respectively, but these returns increase in the comparison periods +60 days, +30 days, and +20 days (0.04%, 1.08%, and 1.30%) but slightly decrease in +10 days (-1.19). Despite the decrease of returns after the maintaining dividend announcement in the pre-crisis sample, the post-crisis sample produced unexpected results except ±10 days period (Table 2).

Figure 3: Bad News/Dividend Omissions: Pre-crisis Sample (1988-1997)

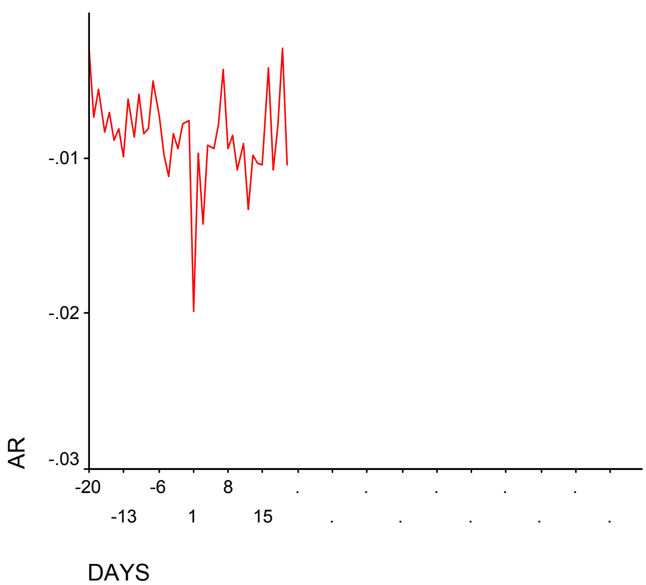
± 60 Days of the Event Day



± 30 Days of the Event Day



± 20 Days of the Event Day



± 10 Days of the Event Day

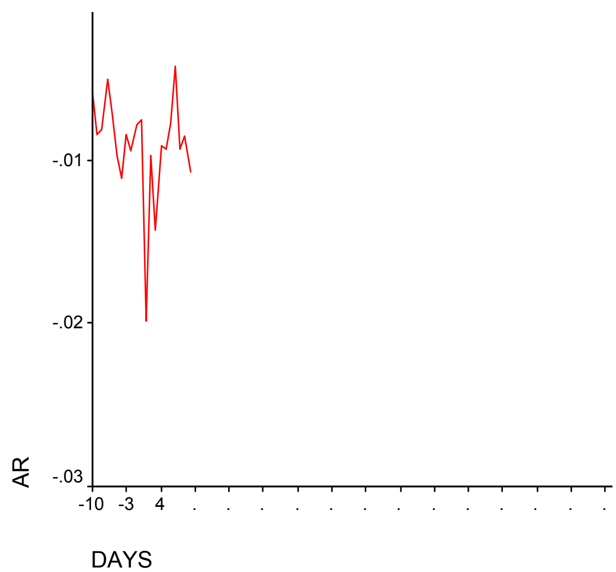
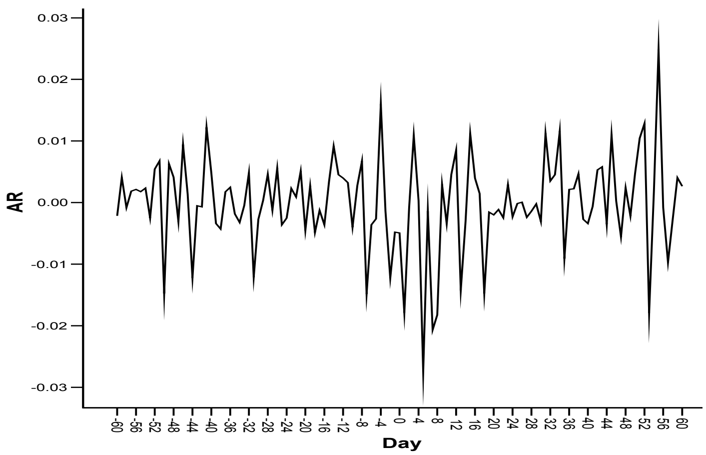
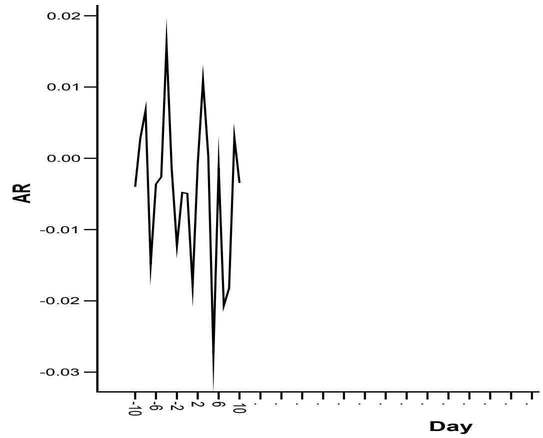


Figure 4: Bad News/Dividend Omissions: Post-crisis Sample (1999-2003)

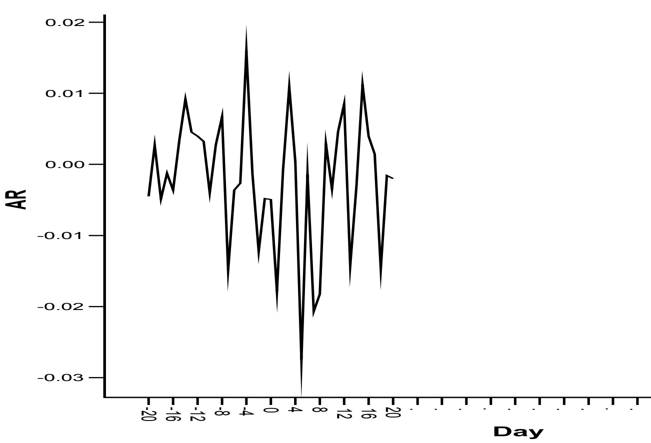
± 60 Days of the Event Day



± 10 Days of the Event Day



± 20 Days of the Event Day



± 30 Days of the Event Day

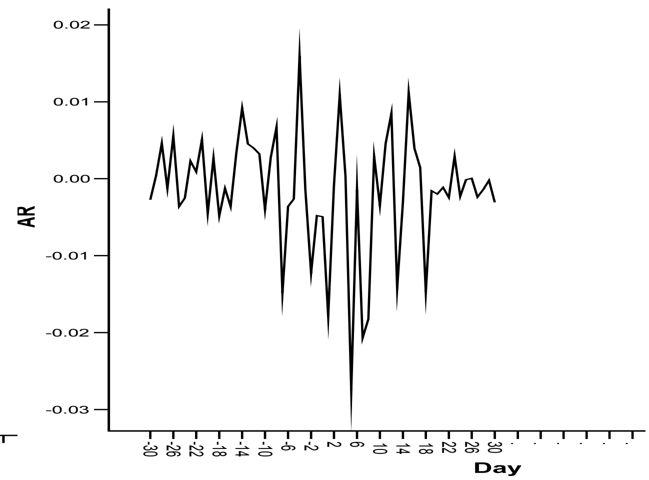


Table 2: Paired Samples Statistics

Good News						
Period	**Pre-crisis Period (1988-1997)**			**Post-crisis Period (1999-2003)**		
	N	Mean	Std. Deviation	N	Mean	Std. Deviation
- 60 Days	198	-0.0001	0.0035	70	-0.0020	0.0040
+ 60 Days	198	-0.0012	0.0029	70	-0.0020	0.0065
- 30 Days	198	-0.0003	0.0021	70	-0.0022	0.0046
+ 30 Days	198	-0.0015	0.0035	70	-0.0023	0.0059
- 20 Days	198	-0.0001	0.0022	70	-0.0028	0.0046
+ 20 Days	198	-0.0015	0.0036	70	-0.0024	0.0071
- 10 Days	198	-0.0001	0.0020	70	-0.0031	0.0039
+ 10 Days	198	-0.0001	0.0036	70	-0.0028	0.0067
	Bad News					
Period	**Pre-crisis Period (1988-1997)**			**Post-crisis Period (1999-2003)**		
	N	Mean	Std. Deviation	N	Mean	Std. Deviation
- 60 Days	79	0.0075	0.1036	46	0.0002	0.0060
+ 60 Days	79	-0.009	0.0034	46	-0.0004	0.0090
- 30 Days	79	-0.0076	0.0019	46	0.0001	0.0061
+ 30 Days	79	-0.0089	0.0035	46	-0.0031	0.0091
- 20 Days	79	-0.0077	0.0019	46	-0.0003	0.0071
+ 20 Days	79	-0.0096	0.0037	46	-0.0041	0.0110
- 10 Days	79	-0.0081	0.0018	46	-0.0018	0.0089
+ 10 Days	79	-0.0103	0.0042	46	-0.0076	0.0125
No News						
Period	**Pre-crisis Period (1988-1997)**			**Post-crisis Period (1999-2003)**		
	N	Mean	Std. Deviation	N	Mean	Std. Deviation
- 60 Days	75	0.0004	0.0033	42	-0.0055	0.0506
+ 60 Days	75	-0.0005	0.0035	42	0.0004	0.0814
- 30 Days	75	0.0005	0.0043	42	-0.0115	0.0023
+ 30 Days	75	-0.0016	0.0030	42	0.0108	0.1140
- 20 Days	75	0.0006	0.0050	42	-0.0122	0.0020
+ 20 Days	75	-0.0020	0.0028	42	0.0130	0.1282
- 10 Days	75	0.0010	0.0061	42	-0.0116	0.0018
+ 10 Days	75	-0.0021	0.0035	42	-0.0119	0.0099

The correlation coefficients between abnormal returns of observation periods and comparison periods of the pre-crisis sample are .039, -.219, -.037, and -.407 and the probability values are .770, .244, .876, and .243 respectively for ±60 days, ±30 days, ±20 days and ±10 days. However, the correlation coefficients between abnormal returns of observation periods and comparison periods of the post-crisis sample are -.009, .139, .461, and -.234 and the probability values are .947, .464, .041, and .516 respectively for ±60 days, ±30 days, ±20 days and ±10 days. Though the correlation coefficients between the abnormal returns of the observation and comparison periods for dividend maintaining announcements in all the study sample (±60 days, ±30 days, ±20 days, and ±10 days) are in the opposite direction both in the pre and post crisis, the results are not statistically significant at the higher level in either pair (Table 3).

Table 3: Good News: Paired Samples Correlation

Good News				
Period	Pre-crisis Period (1988-1997)		Post-crisis Period (1999-2003)	
	N	Correlation	N	Correlation
± 60 Days	198	-0.232*	70	-0.019
± 30 Days	198	-0.076	70	-0.062
± 20 Days	198	-0.243	70	-0.228
± 10 Days	198	0.116	70	-0.108
Bad News				
Period	Pre-crisis Period (1988-1997)		Post-crisis Period (1999-2003)	
	N	Correlation	N	Correlation
± 60 Days	79	0.029	46	-0.200
± 30 Days	79	-0.330*	46	-0.018
± 20 Days	79	-0.603***	46	-0.134
± 10 Days	79	-0.630*	46	-0.222
No News				
Period	Pre-crisis Period (1988-1997)		Post-crisis Period (1999-2003)	
	N	Correlation	N	Correlation
± 60 Days	75	0.039	42	-0.009
± 30 Days	75	-0.219	42	0.139
± 20 Days	75	-0.037	42	0.461**
± 10 Days	75	-0.407	42	-0.234

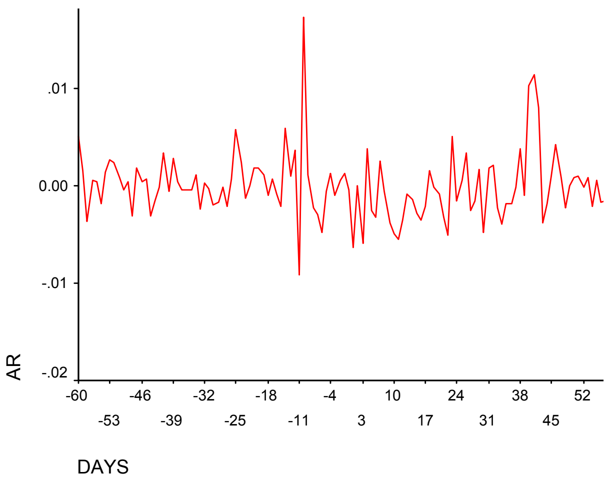
****Significant at 1% level **Significant at 5% level *Significant at 10% level*

The mean difference between the abnormal returns of the observation and the comparison periods of the pre-crisis sample are .0009, .0021, .0026, and .0031 respectively for ±60 days, ±30 days, ±20 days and ±10 days. However, the mean difference between the abnormal returns of the observation and the comparison periods of the post-crisis sample are -.0059, -.0224, -.0252, and .0003 respectively for ±60 days, ±30 days, ±20 days and ±10 days. The t-values and the probability values of the pre-crisis sample are 1.491, 2.005, 1.974, and 1.207, and .141, .054, .063, and .258 respectively for ±60 days, ±30 days, ±20 days and ±10 days. However, the t-values and the probability values of the post-crisis sample are -.477, -1.077, -.885, and .100, and .635, .291, .387, and .923 respectively for ±60 days, ±30 days, ±20 days and ±10 days. These results imply that the mean difference of the abnormal returns between the observation and comparison periods is not significantly different from zero. Nevertheless, the sequence charts of the abnormal returns for the event study periods (Figures 5 and 6) also support the empirical evidence of this study. Therefore, the empirical evidence of this study contradicts with the previous studies that security prices do not react to the dividend maintaining announcements (see Table 4).

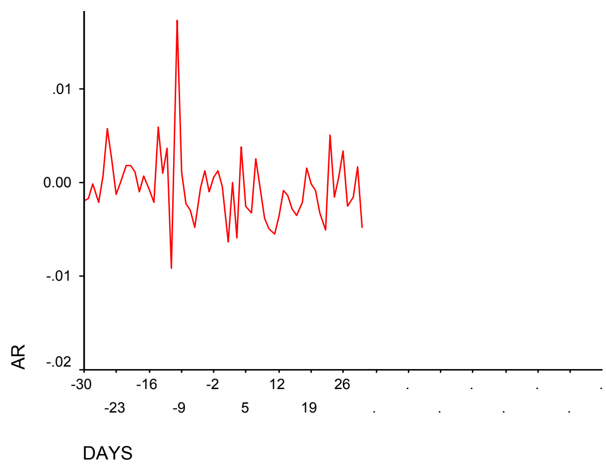
Despite a slight change in the post-crisis sample, *o*verall the empirical results failed to reject the announcement effect hypothesis that the security returns in Bangladesh stock market decrease after dividend initiations, omissions, and dividend maintenance in the pre-crisis sample but scenario is little bit different in the post-crisis sample, i.e., security returns increase in dividend initiations and maintenance but decrease in dividend omissions. Nevertheless, the signaling effect of the announcements appears ineffective as because t-statistics are not significant at a very high level. This is the clear symptom of ineffectiveness of dividend announcements in the emerging market of Bangladesh. Therefore, the announcement of dividends does not carry any new information to the market. These results also strongly reject the signaling theory of dividends. Most important reasons for the ineffectiveness of the announcements of dividend in an emerging market are the insider trading and because of that, information incorporates the market prices before the announcements. The other reason is that the insiders are involved in motivated trading before and after the announcement of dividends. As we already mentioned that insiders hold higher percentage of stocks in Bangladesh. Usually insiders start to buy back shares before the annual general meeting (AGM) for higher voting rights that causes higher demand of shares and consequently higher share returns and as insiders off load shares after AGM, that causes huge supply of shares and consequently returns decrease after the annual general meeting.

Figure 5: No News/Dividend Maintenance: Pre-crisis Sample (1988-1997)

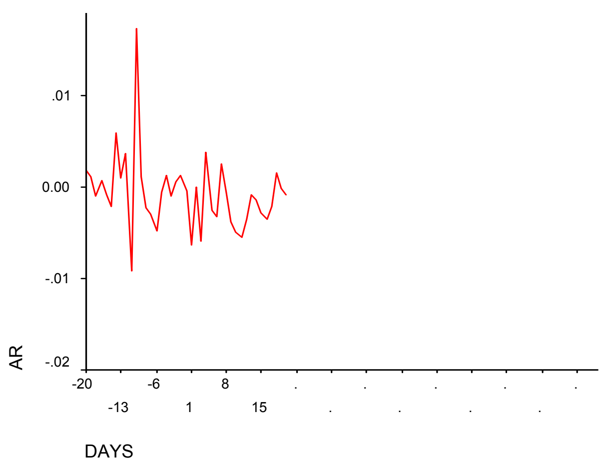
± 60 Days of the Event Day



± 30 Days of the Event Day



± 20 Days of the Event Day



± 10 Days of the Event Day

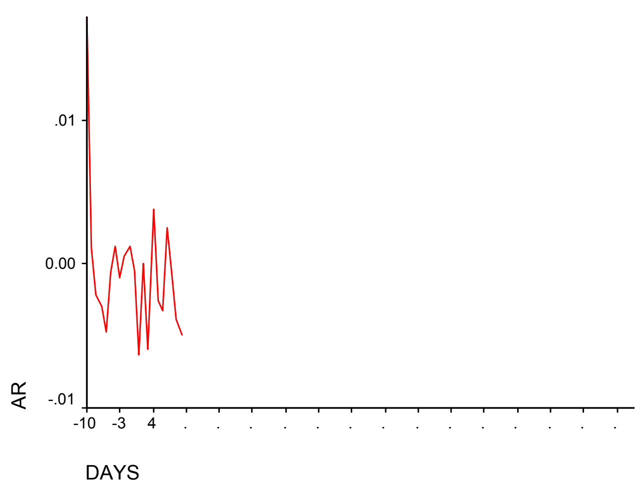


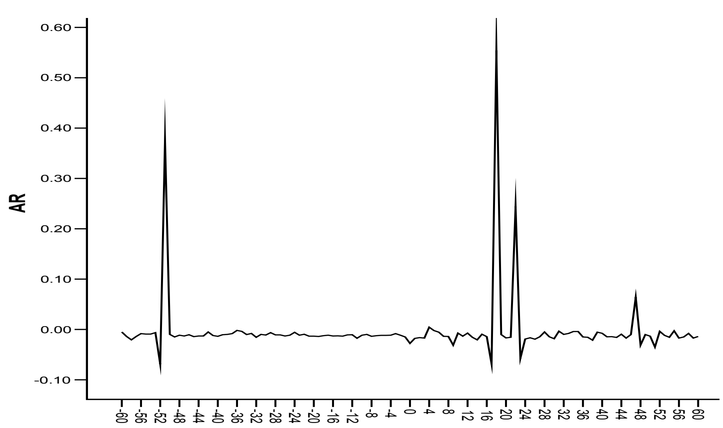
Table 4: Paired Samples T-Test

Good News										
Period	**Pre-crisis Period (1988-1997)**					**Post-crisis Period (1999-03)**				
	Mean	Std. Deviation	95% Confidence Interval of the Difference		T Value	Mean	Std. Deviation	95% Confidence Interval of the Difference		T Value
			Lower	Upper				Upper	Lower	
± 60 Days	0.0012	0.0050	-0.0001	0.0025	1.772*	0.0001	0.0077	-0.0020	0.0020	0.023
± 30 Days	0.0013	0.0042	-0.0003	0.0029	1.681	0.0001	0.0077	-0.0028	0.0030	0.086
± 20 Days	0.0014	0.0047	-0.0008	0.0036	1.340	-0.0004	0.0092	-0.0048	0.0040	-0.208
± 10 Days	0.0001	0.0039	-0.0028	0.0029	0.016	-0.0003	0.0080	-0.0060	0.0055	-0.105
Bad News										
Period	**Pre-crisis Period (1988-1997)**					**Post-crisis Period (1999-03)**				
	Mean	Std. Deviation	95% Confidence Interval of the Difference		T Value	Mean	Std. Deviation	95% Confidence Interval of the Difference		T Value
			Lower	Upper				Lower	Upper	
± 60 Days	0.0166	0.1035	-0.0101	0.04.34	1.243	0.0006	0.0117	-0.0024	0.0036	0.410
± 30 Days	0.0013	0.0045	-0.0004	0.0029	1.559	0.0032	0.0110	-0.0009	0.0073	1.577
± 20 Days	0.0019	0.0050	-0.0004	0.0043	1.683	0.0038	0.0139	-0.0027	0.0103	1.230
± 10 Days	0.0022	0.0055	-0.0017	0.0061	1.261	0.0058	0.0168	-0.0063	0.0178	1.084
No News										
Period	**Pre-crisis Period (1988-1997)**					**Post-crisis Period (1999-03)**				
	Mean	Std. Deviation	95% Confidence Interval of the Difference		T Value	Mean	Std. Deviation	95% Confidence Interval of the Difference		T Value
			Lower	Upper				Lower	Upper	
± 60 Days	0.0009	0.0048	-0.0003	0.0021	1.491	-0.0059	0.0962	-0.0308	0.0189	-0.477
± 30 Days	0.0021	0.0058	-0.0001	0.0042	2.005*	-0.0224	0.1137	-0.0648	0.0201	-1.077
± 20 Days	0.0026	0.0058	-0.0002	0.0053	1.974*	-0.0252	0.1273	-0.0848	0.0344	-0.885
± 10 Days	0.0031	0.0081	-0.0027	0.0089	1.207	0.0003	0.0105	-0.0072	0.0079	0.100

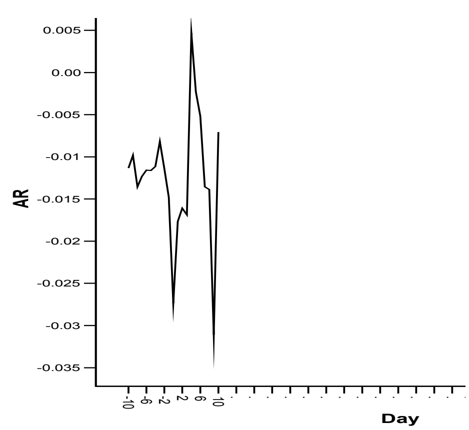
**Significant at 10% level*

Figure 6: No News/Dividend Maintenance: Post-crisis Sample (1999-2003)

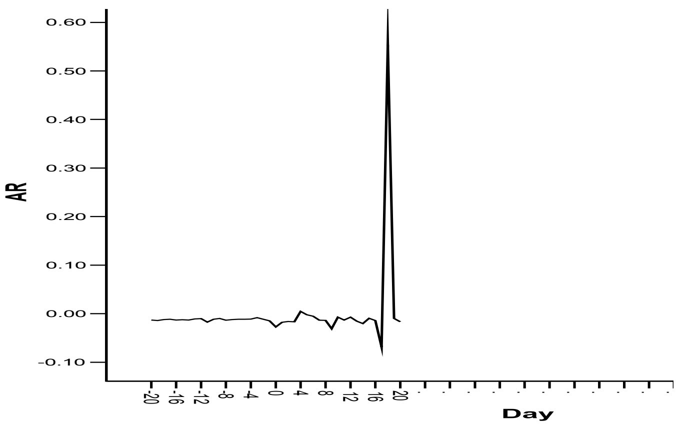
± 60 Days of the Event Day



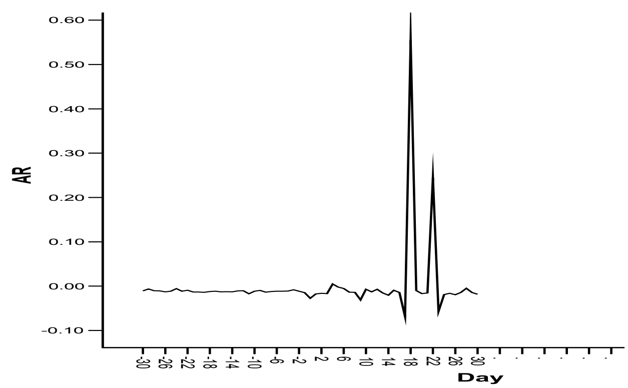
± 10 Days of the Event Day



± 20 Days of the Event Day



± 30 Days of the Event Day



Ineffectiveness of dividend announcements also causes for many other reasons including companies announce dividends but they often do delay in paying dividends to shareholders, after the book closure companies take long time to transfer the ownership, etc. For these and many other reasons, the shareholders are always skeptical about the activities of the management and they do not trust management with full confidence. Finally, the lower level of law enforcement in the market and ineffectiveness of the regulatory bodies is also a significant cause of distortion in the market.

CONCLUSION

A vast majority of the studies found dividend announcements as a strong signaling device, which influence the security prices but the issue of the effect of dividend announcements on security prices is still inconclusive. The major objective of this paper is to identify whether dividend announcements convey information to the market or whether investors consider the announcement of dividends as a signal of the firm's future prospects, i.e., to see the security price reaction to the announcement of dividends in an emerging market. The empirical results reject the dividend-signaling hypothesis that dividend announcements do not convey any information about the companies listed on the Dhaka Stock Exchange. After the financial crisis in Dhaka stock market of Bangladesh in 1998, there were significant changes in institutional setting but there was no change in the legal framework as the controlling mechanism. The market also fails to come up with a significant reform following the financial crisis in Bangladesh, therefore, the reform does not help to improve the market scenario.

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DOES CORRUPTION MATTER FOR NIGERIA LONG RUN GROWTH: EVIDENCE FROM COINTEGRATION ANALYSES AND CAUSALITY TESTS?

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ABSTRACT

The study examines the relationship between corruption and economic growth in the Nigeria economy for sample periods ranging from 1970 to 2004. Johansen's maximum likelihood cointegration techniques and Granger causality tests were applied to annual, national-level data. The results of this study indicate that corruption is cointegrated with Economic growth in Nigeria. In addition, for Nigeria, the study found a one-way causality from corruption to economic growth. These findings provide a statistical confirmation of unfavorable effects of corruption on economic performance as widely hypothesized in economic literature. For policy, the results of this study suggest that the current anti corruption drive in the country should be more vigorously pursued as this result indicated that it has important consequence on economic growth aspirations of the country.

INTRODUCTION

Corruption is a modern scourge that has become endemic in many societies, both developed and underdeveloped. Corruption is one of the most harmful phenomena of the age, its devastating social consequences threatening the safety and security of peoples, states and democratic institutions. This becomes more worrisome for most countries of Sub Saharan African (SSA) with reputation for enshrined institutionalized corruption that has become widespread to the extent that it is no longer considered an exception but the rule. The Nigeria economy has a reputation of being corrupt. Transparency International (TI) rated Nigeria the world's third most corrupt nation; it as well accused the country of being on the lead of African nations allegedly slowing down the fight against corruption in the continent. Raymond Baker, in a paper Money Laundering and Flight Capital: The Impact on Private Banking, wrote: "the biggest single thief in the world in the 1990s was almost certainly the late military dictator, Sane Abacha, with $12 to $16 billion passing out of Nigeria in corrupt and tax evading money during his murderous five year regime." Startling revelations of corrupt practices by key operators at all levels and arms of current Nigeria's fledging democracy point to the fact that corruption still poses serious threat to national survival.

The purpose of this paper is to seek an empirical understanding and an addition to current literature on the association between corruption and economic growth by using cointegration analyses to examine the relationship between these phenomenons in Nigeria. This method is quite common in the area of economics and finance in general, but has not been applied extensively to the study of corruption. In addition, causality tests will be performed to the Nigerian economy data to examine a causal link between corruption and economic growth.

The rest of the paper is organised as follows: Section two provides an outline of theoretical linkages between corruption and economic growth. It also contains a review of relevant empirical literature in the area. Section three examines definitional and measurement issues in empirical analyses of corruption. An operational measure of capital flight as a proxy for corruption in the Nigerian economy used in the study is also described in section three. Section four set out the methodological framework for the study. The framework established in section four is subjected to econometric analysis in section five and section six concluded with summary and policy implications of the main findings.

THEORETICAL AND EMPIRICAL LITERATURE

Much of the concern over corruption in recent years has dealt with its effects upon economic development; hence, literature is replete with theoretical and empirical evidences on the relationship between corruption and economic growth. The general conclusion is that corruption slows down the long-term growth of an economy through a wide range of channels. Corruption is both a tax and source of uncertainty to investment decisions, and both diminish incentives to invest. Businesspersons interpret corruption as a species of tax. This is because they are often aware that a bribe is required before an enterprise can be started and, in addition, corrupt officials may also lay claims to part of the proceeds from the investment. In addition, since corruption is shrouded in secrecy, they face the uncertainty that the corrupt official will not fulfill his part of the bargain. Mauro (1995) presents some strong empirical evidence to help prove the negative relationship between corruption and long-term growth. Wei (1997) argues that corruption is much more costly than ordinary taxes because it generates uncertainty in addition to the tax burden.

Corruption is a form of seeking economic rents by creating artificial limitations. Every day private firms spend vast amounts of money attempting to convince legislators to grant monopolies or otherwise restrict competition so that some industry or individual can realize a rent. Throughout the world bureaucrats and people in authority are indefatigably maneuvering to position themselves in a tiny monopoly where they can be bribed for issuing a license, approving an expenditure, or allowing a shipment across a border. Studies have shown that these rent-seeking activities exact a heavy economic and social toll. Since rent seeking is often more lucrative than productive work, talents will be misallocated. Financial incentives may lure the more talented and better educated to engage in rent seeking rather than productive work, which in turn results in adverse consequences for the country's economic growth. Ehrlich and Lui (1999) present a balanced growth model to show that in some equilibrium officials spend a substantial amount of time and effort in seeking and accumulating political capital, which is not socially productive.

One specific channel through which corruption may harm economic performance is by distorting the composition of government expenditure. Corrupt politicians may be expected to spend more public resources on those items on which it is easier to exact large bribes and keep them secret, for example, items produced in markets where the degree of competition is low and items whose value is difficult to monitor. Corrupt politicians might therefore be more inclined to spend on fighter aircraft and large-scale investment projects than on textbooks and teachers' salaries, even though the latter may promote economic growth largely than the former. Mauro (1998) concludes that corruption affects the composition of government expenditure. When corruption is serious, there is much less government expenditure on education than on large infrastructure and defense projects. In addition, Mauro finds that corruption also lowers the quality of infrastructure projects and pubic services.

Definitional and Measurement Issues

Corruption is an illicit activity and hence is difficult to define, conceptualize, and measure. Corruption is defined in the literature in a variety of ways. When viewed from the perspective of public interest, corruption is defined as "the abuse of public power for private gain." From the political-economy angle, it is described as "charging of a price for the provision of a public service in excess of the official tariff." When corruption is viewed as a behavioral phenomenon, it is defined as engaging in activities deemed illegal by society. Thus it becomes difficult to reduce the phenomenon to a single definition as it covers such a vast and growing assortment of activities which may include, but are not limited to, drugs trafficking, money laundering, corporate crime, arms sales, sex trade and abuse, art and antique fraud, human trafficking, capital flight, etc. The only common denominator linking these activities together is that they are all undertaken outside the legality of the system.

Corruption is a classic example of an observable phenomenon that is not quantifiable since there cannot be statistics on a phenomenon, which by its very nature is concealed. Corruption is usually a clandestine activity and frequently does not have a direct victim. Moreover, those with knowledge of a corrupt act generally have an interest in concealing it. A related complication, as argued by Rose-Ackerman (1999) and Lamsdoff (1999), has to do with what is meant by "levels" of corruption. Is it the number of corrupt actions over time, the size of the stakes involved or the level of government at which they occur? For these reason corruption remains impossible to measure directly.

Attempts to measure corruption have involved several approaches. These include perception-based corruption indices, which draw upon opinion surveys and expert estimates of how corrupt various countries are. This index is widely associated with the Transparency International's Corruption Perception Index (Transparency International, 2001). Schlessinger and Meir (2002) made comparisons based upon arrest or conviction data for corruption offences. PriceWaterhouseCoopers (2001), provides an hybrid approach, known as the "Opacity Index", which incorporates a multi-component indices perceptual ratings of hard economic and social data and indicators such as interest premiums paid on sovereign debt by nations with transparency problems.

While these perceptual scales have helped in understanding the cause and effects of corruption, as well as put pressure on governments and society to address the corruption problem, they are obviously imperfect owing to their subjective nature. Perception data are just estimates of how corrupt a society are thought to be, and are thus open to influence and distortion from full range of factors affecting any human judgments. The more obvious limitation of these data sets is that it treats whole societies as units of analysis, and thus it becomes far removed from analysis that have specific country or provinces as focus. There are no obvious answers to these problems, yet finding useful ways to compare level of corruption over time remains a necessity for an empirical focused study such as the present study that need good data for building models and testing hypotheses.

<u>The Measure of Corruption Used in This Study</u>

To address the problems highlighted in the previous section, this paper adopts the residual method estimates of capital flight as a proxy for corruption in Nigeria. Klitgaard, Maclean-Abaroa and Parris (2000) contend that the most useful approach in measuring corruption is to track changes in aspects of governance that create incentives for corruption, or reveal its effects or both. One such incentive in Africa is the globalization of markets and developmental inequalities that have engendered cross-border corruption involving international interests, actors, capital and economic processes. Ample empirical evidence has demonstrated that incentives for cross-border corruption in form of illegal external capital / asset accumulation, or what economists term capital flight is associated with large-scale corruption among African leaders. Bardhan, (1997) opined that it is plausible that certain types of capital flight, defined as transfer of money abroad, usually in US dollars, often under questionable circumstances are functions of the government, or worse still governmental corruption.

Boyce and Ndikumana (2002) noted that in most African countries, instead of financing investment or consumption with public funds, a substantial fraction was captured by African political elites and channeled abroad in the form of capital flight. Through this process public funds, as well as public external debts (contracted via borrowing by African government or by private firms with government guarantees) were transformed into private external assets. Activities and actions of political office holders, especially the military elites, in Nigeria in the last three decades perfectly fits in the mechanisms by which resources are channeled abroad as capital flight. These include embezzlement of borrowed funds, kickbacks on government contracts, trade mis-invoicing, misappropriation of revenues from state-owned enterprises and smuggling of natural resources. As the government they headed incurred large external debts, a number of individual military rulers amassed large personal fortunes. A substantial part of which were held abroad. For instance, the Swiss bank accounts of the family of General Sani Abacha, who ruled Nigeria for five years, reportedly contain as much as $2billion US dollars at the time it was frozen in 1999 (Onishi, 1999). In addition, a US Senate enquiry in the same year revealed that the Abacha family also held multi-million dollar accounts with Citibank in London and New York (Gerth, 1999; O'Brien, 1999).

The residual method measures capital flight indirectly by comparing the sources of capital inflows (i.e. net increases in external debt and the net inflow of foreign investment) with the uses of these inflows (i.e., the current account deficit and additions to foreign reserves). This approach starts from the standard balance of payments framework. In principle, if the balance of payments statistics were to be used (reported by the International Monetary Fund Balance of Payments Statistics), the uses and sources of funds should be equal. However, since these statistics may not accurately measure flows, and in particular private capital flows, World Bank statistics on the change in the external debt are used instead. If the sources, calculated by using World Bank debt data, exceed the uses of capital inflows, the difference is termed as capital flight. The residual method acknowledges the difficulties of separating abnormal from normal capital outflows and, therefore, measures all unrecorded private capital outflows as being capital flight.

According to the residual method, capital flight is calculated as follows:

$$KF_r =)ED + FI - CAD -)FR \qquad (1)$$

where KF_r is capital flight according to the residual method, *)* denotes change, *ED* is stock of gross external debt reported in the World Bank data, *FI* is the net foreign investment inflows, *CAD* is the current account deficit and *FR* is the stock of official foreign reserves. Annual data series on these variables for the period 1970 to 2004 are sourced from International Financial Statistics (IFS) publications of the International Monetary Fund (IMF)

METHODOLOGY, DATA AND ESTIMATION TECHNIQUES

We set out a simple model to test for the existence of any long-run relationship and potential causality between corruption and GDP growth rate. Equation 2 specifies our simple model:

$$\text{LRGDP} = \alpha + \beta \text{LCORt} + \text{Ut} \qquad (2)$$

Where LRGDP denotes the logarithm of real GDP, α and β are estimated constants and LCOR is the logarithms of the *capital flight* proxy for corruption estimated from equation (1). For modeling purposes, the variables are in natural logarithms; thus, the first differences can be interpreted as the rate of growth.

The study analyzes the link between corruption and economic growth in a time series data of Nigeria for the 1970-2004 periods, via the cointegration tests and granger causality analyses. The data on corruption is proxy by an estimate of capital flight flows from Nigeria. The data, sources and estimates of capital flight are as presented in section 3.1. Economic growth is proxy by the percentage rate of growth of real gross domestic product. Data on this variable is sourced from Statistical Bulletin publications of the Central Bank of Nigeria.

Co-integration analysis provides potential information about the long-term equilibrium relationship of the model. It is now widely recognized following Granger and Newbold (1974) that most economic series exhibit a non-stationary (unit-root) pattern in their levels, i.e. the means and variances are time dependent and such variables are said to be I(1) (Holdenand Perman, 1994). The implication is that all computed statistics in a regression model, which use these means and variances are also time dependent. It implies that such variables fail to converge to their true values as the sample size increases (Rao, 1994). If, after differencing, the variables become stationary then they are referred to as being I(0). The technique of co-integration is not only essential, but also necessary in estimating an equilibrium relationship with unit root or non-stationary variables to determine the presence of a long-run relationship.

Appropriate tests to determine whether a time series is integrated of order one against the alternative of zero order integration include those developed by Fuller (1976), Dickey and Fuller (1981), Phillips (1987), and Perron (1988) and others. In addition, there are various approaches to estimating cointegrating regressions. Two broad approaches are available: (a) Engle-Granger (1987), (b) Johansen (1988), and Johansen-Juselius (1990). The approach developed by Johansen (1988) and Johansen-Juselius (1990), which is based on the full information Johansen Maximum Likelihood method (JML) is preferred and used in this study. The first approach is popular due to its simplicity and ease of calculation. However, there are some problems with the Engle and Granger (1987) procedure. With the Engle-Granger approach, the estimation of the long-run equilibrium regression requires that the researcher place' one variable on the left-hand side and use the other as regressors. For example, in the case of two variables, it is possible to run the Engle-Granger test for co-integration by using the residuals from either of the following two 'equilibrium' regressions:

$$y_t + \beta_{10} + \beta_{11} z_t + e_{1t} \text{ and } z_t + \beta_{20} + \beta_{21} y_t + e_{2t}.$$

However, a problem arises from finite samples. As the sample size increases, asymptotic theory indicates that the test for a unit root in the first error sequence becomes equivalent to the test for a unit root in the second error sequence (Enders, 1995). Furthermore, it is possible to find that one regression indicates the variables are co-integrated, whereas reversing the order indicates no cointegration (Enders, 1995). The VAR approach considers this possibility and treats all variables as potentially endogenous. Moreover, recent Monte Carlo evidence strongly favors the Johansen Maximum Likelihood method (JML) approach over the Engle-Granger's (Dejong, 1992). The Hypothesis is that "If variables are co-integrated, they share a long-run relationship and will move closely together over time."

Johansen's approach is to estimate the Vector Error Correction Mechanism (VECM) by maximum likelihood, under various assumptions about the trend or intercept parameters and the number r of cointegrating vectors, and then conduct likelihood ratio tests. Assuming that the VECM errors U_t are independent $N_k[0,S]$ distribution, and given the cointegrating restrictions on the trend or intercept parameters, the maximum likelihood $L_{max}(r)$ is a function of the cointegration rank r. Johansen proposes two types of tests for r:

> *The lamda-max test:* This test is based on the log-likelihood ratio $\ln[L_{max}(r)/L_{max}(r+1)]$, and is conducted sequentially for $r = 0,1,..,k-1$. The name comes from the fact that the test statistic involved is a maximum generalized eigenvalue. This test tests the null hypothesis that the cointegration rank is equal to r against the alternative that the cointegration rank is equal to $r+1$.
>
> *The trace test:* This test is based on the log-likelihood ratio $\ln[L_{max}(r)/L_{max}(k)]$, and is conducted sequentially for $r = k-1,...,1,0$. The name comes from the fact that the test statistic involved is the trace (= the sum of the diagonal elements) of a diagonal matrix of generalized eigenvalues. This

test tests the null hypothesis that the cointegration rank is equal to r against the alternative that the cointegration rank is k. The latter implies that X_t is trend stationary.

Both tests have non-standard asymptotic null distributions. Moreover, given the cointegration rank r Johansen also derives likelihood ratio tests of the cointegrating restrictions on the intercept or trend parameters.

In addition to cointegration analyses, this study also conducts the Granger causality tests of the same variables to detect any causal link between economic performance and corruption. Traditionally, causality tests between two stationary series are based on Granger's (1969) definition for causality. Formally, series y_t "Granger-causes" series x_t if series x_t can be predicted better by using past values of series y_t than by using only the historical values of series x_t. In other words, y_t fails to Granger-cause x_t if, for all s > 0, the conditional probability distribution of x_{t+s} given (x_t, x_{t-1},...) is the same as the conditional probability distribution of x_{t+s} given both (x_t, x_{t-1},...) and (y_t, y_{t-1},...). That is, y_t does not Granger-cause x_t if : $Pr(x_{t+s}|\boldsymbol{X}_{t,-l}) = Pr(x_{t+s}|\boldsymbol{X}_{t,-l},\boldsymbol{Y}_{t,-l})$ where $Pr(\cdot)$ denotes conditional probability, $\boldsymbol{X}_{t,-l}$? (x_t, x_{t-1},..., x_{t-l}) and $\boldsymbol{Y}_{t,-l}$? (y_t, y_{t-1},..., y_{t-l}). Granger (1969) proposes the test for causality between x_t and y_t by running a set of regressions:

$$Xt = \theta 0 + \overset{n}{\underset{i=1}{\theta}} a i x t{-}i + \overset{n}{\underset{i=1}{\theta}} b i y t{-}i + u t \qquad 3$$

$$y t = \theta 1 + \overset{n}{\underset{i=1}{\theta}} a i x t{-}i + \overset{n}{\underset{i=1}{\theta}} d i y t{-}i + v t \qquad 4$$

where a_0 and a_1 are constants, a_i, b_i, c_i, and d_i are parameters, and u_t and v_t are uncorrelated error terms with zero means and finite variances. The null hypothesis that y_t (x_t) does not Granger-cause x_t (y_t) is rejected if the b_i (c_i) coefficients are jointly significantly different from zero, using a standard F test. Bi-directional causality (or feedback) exists if both the b_i and c_i coefficients are jointly different from zero.

EMPIRICAL ESTIMATION AND INTERPRETATION OF RESULTS

Unit Roots Tests

Before we estimate the equation, we determine the underlying properties of process that generate our time series variables, that is, a test of the stationary properties of the variables. Macroeconomic data often appear to possess stochastic trend that can be removed by differencing the variables. We use the Augumented Dickey Fuller (ADF) t-test for testing the order of integration. Assuming there is no trend, the ADF test can be formulated as follows:

$$\Delta y_t = \delta . y_{t-1} + \sum_{i=1}^{k} \delta . \Delta y_{t-1} + e_t \qquad 5$$

The null hypothesis being tested is that $* = 0$ (random walk with a drift) against the alternative of stationarity. The results are as presented in Table 1. As shown in the table, the ADF test indicated that our series are non-stationary at their levels. However, all the first differenced series turn out to be stationary at the 5% level of significance, the critical value computed by McKinnon is -3.02. All the first differenced test results have t-statistics exceeding McKinnon's critical value, so that the hypothesis $* = 0$ could now be rejected.

Table 1: Augmented Dickey Fuller Test of Unit Roots

Variables	**ADF Test Statistics**
A. Series in Levels	
RGDP	-2.5734
COR	-2.7548
B. Series in First Differences	
RGDP	-3.7424*
COR	-4.0426*

*ADF is the Augmented Dickey- Fuller test; it gives the t-statistics from a specification that includes a constant, trend and two (2) lagged changes in the dependent variable. A * indicates rejection of the null hypothesis (□ = 0) of non-stationarity at the 5% level of significance. MacKinnon critical value for rejection of a unit root for ADF at 5% is -3.02.*

Tests for Cointegration

We employed the Johansen Cointegration test to check for cointegration among the time series. This become necessary because our variables contain unit roots in the level, cointegration is the appropriate dynamic modeling technique for them. A linear combination of these variables is identified such that this combination is stationary. If such combination exists, then the variables are said to be cointegrated. If variables are co-integrated, they share a long-run relationship and will move closely together over time. This means that the difference between such variables was stable over time and there is some degree of convergence in the long run. The estimation of a VAR model requires the explicit choice of lag length in the equation of the model. Following Judge et.al (1988), Akaike's AIC criterion was used to determine the lag length of the VAR model. The chosen lag length is one that minimized the following: $AIC_{(n)} = \ln\det\Sigma_n + \{2d^2n\}/T$…where d is number of variables, T is the sample size, and Σ_n is the estimate of the residuals of the variance-covariance matrix obtained with a VAR. The model that minimized AIC turns out to be the one with 2 lag lengths. It is hypothesized that there exists a long-run relationship between real GDP and corruption proxied with the residual estimate of capital flight from the Nigerian economy.

The results of the test are as shown in Table 2 below. Panel A reports the so-called trace statistics, while Panel B reports the maximal eigenvalue statistics. The first column shows the number of cointegration relations under the null hypothesis, the second is the ordered eigenvalue of the *II* matrix, the third column is the trace statistic, and the last columns are the 5% and 1% critical values. It should be noted that the (nonstandard) critical values are taken from Osterwald-Lenum (1992), which differ slightly from those reported in Johansen and Juselius (1990). The trace statistic tests the null hypothesis of r cointegrating relations against the alternative of k cointegrating relations, where k is the number of endogenous variables, for R = 0, 1,….., K – 1. Using these trace statistic, we test for the number of co-integrating relationship between LRGDP and LCOR. Given that we only have two variables, we expect that at least one cointegrating vector is present.

To test the null hypothesis r =zero against the general alternative r =1, or 2 we use the λ-trace statistic and the Eigen value. Since the null hypothesis is r=0 and there are two variables (i.e. n =2), the summation in the estimated equations runs from 1 to 2. The calculated value for the trace statistics is 19.95054 and comparing this calculated values to the critical values provided by Johansen and Juselius (1990), the null hypothesis of cointegration can be accepted at both 5% and 1% critical levels. Thus, at the 90% level, the restriction is binding and we conclude that the variables are cointegrated.

Table 2: Unrestricted Cointegration Rank Test

Panel A: Trace Statistics				
Hypothesized No. of CE(s)	Eigenvalue	Trace Statistics	5 Percent Critical Value	1 Percent Critical Value
None*	0.384686	19.95054	15.41	20.04
At Most 1 *	0.183172	19.95054	3.76	6.65
Panel B: Maximal Eigenvalue Statistics				
Hypothesized No. of CE(s)	Eigenvalue	Trace Statistics	5 Percent Critical Value	1 Percent Critical Value
None*	0.384686	19.95054	14.07	18.63
At Most 1 *	0.183172	5.867474	3.76	6.65

**(**) denotes rejection of the hypothesis at 5% (1%) significance level. Trace statistics indicates 2 cointegrating equation(s) at 5% (1%) significance level. Max-Eigen Statistic indicates 2 cointegrating equation(s) at 5% significance level*

Since the trace test has a very general alternative hypothesis, we need to test a more specific hypothesis. To do so, we apply the lambda-maximal test. Here we test the null hypothesis r=0 against the specific alternative of r=1. The calculated and critical values for n-r =2 is as shown in the table. The null hypothesis of cointegration can still be accepted at both 5% and 1% critical levels.

As reported in Table 2 above, both the trace and the maximal eigenvalues statistics show that the VAR has two cointegrating vectors. This implies that a long run relationship exists among the variables. Hence, the second cointegrating vector is normalized as the growth of real GDP. Thus, the result of the cointegrating vector is given as:

$$RGDP = 2.45 + \underset{(0.261)}{0.18}COR$$

The standard error is indicated in parenthesis. The cointegrating vector indicated that long-run growth of real GDP in Nigeria is negatively and significantly related to levels of corruption. The coefficient suggests that an increase of corruption level by 1 million Naira would reduce economic growth by 18%. This finding provides statistical confirmation of the hypothesized negative impact of corruption on growth in economic literature.

<u>Granger Causality Tests</u>

The Granger causality test requires that all data series involved are stationary. Otherwise, the inference from the *F*-statistic might be spurious because the test will have nonstandard distributions. As shown in Table 1, both the real GDP and corruption series are shown to be I(1). Accordingly, the first-difference series are used to perform the Granger causality tests. The results of the tests are reported in Table 3.

Table 3: Granger Causality Tests Results

Dependent Variable	F-Statistics
Corruption	1.079
Economic growth	5.052*

Reported values are the F-statistics..
** Rejection of the null hypothesis at the 5% level of significance.*

In the corruption equation, there is no causal relationship from economic growth to corruption as the F-statistic was insignificant. In the long run, there is unidirectional causality running from corruption to economic growth as evidenced by statistical significance of F-statistic of the economic growth equation as shown in Table 3 above. Thus, it can be concluded that for Nigeria data, there is uni-directional causality between corruption and economic growth and the causal direction runs from corruption to economic growth.

SUMMARY AND CONCLUSION

A burgeoning empirical literature suggests that the absence of corruption accelerate economic growth, while anecdotal evidences confirmed that corruption is endemic in the Nigeria socio-economic and political polity. This paper therefore seeks an empirical understanding of the association between corruption and economic growth in Nigeria. We achieved this by applying Johansen's maximum likelihood cointegration method and Granger causality test for corruption and economic growth indices from the Nigerian economy, for sample periods from 1970 to 2004.

Both the trace and the maximal eigenvalues statistics show that a long run relationship exists between corruption and economic growth indices in Nigeria. The cointegrating vector indicated that long-run growth of real GDP in Nigeria is negatively and significantly related to levels of corruption. The coefficient suggests that an increase of corruption level by one million Naira would reduce economic growth by 18%. This finding provides statistical confirmation of the hypothesized negative impact of corruption on growth in economic literature.

The Granger causality tests support uni-directional causality between corruption and economic growth and the causal direction runs from corruption to economic growth to further lend credence to the results from cointegration analyses.

These findings provide empirical support for the postulates of negative growth impact of corruption on growth as contained in theoretical and empirical literature. Furthermore, the results of this study appear to suggest that cointegration analysis may be a fruitful way to investigate the issue. For policy, the results of this study suggest that the current anti corruption drive in the country should be pursued very vigorously as this result indicated that it has important consequence on economic growth aspirations of the country. The creation of popular expectations about standards of public service and the right to be free of corruption are important elements of an anti-corruption strategy.

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A SIGNALING MODEL OF CONTROL BLOCK SALES BY ENTREPRENEURS

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ABSTRACT

In this paper, we present a model in which higher-valued managers signal their value by voluntarily submitting to shareholder oversight. If a manager is willing to sell enough stock to release voting control, he is perceived to be of higher quality than if he had defensively maintained control. The implication of the model is that voluntary/control sales by insiders can be good news for the firm. This is consistent with the share-price increases that follow the deaths of entrenched managers.

INTRODUCTION

To some investors, trades by insiders are like tea leaves. As the *Wall Street Journal*'s "street sleuth" recently put it: "Many analysts and investors study the trades of company insiders for cues to buy, hold, or sell shares, believing these individuals or larger shareholders have better insights into a company's prospects and the value of its shares." One popular notion among some of these voyeuristic investors is that insider sales are bad news: insider sales are "ominous" "warning signs" since "it is never encouraging when insiders sell stock," so a smart investor should "jettison any issues where there's been heavy selling."[1] This simplistic trading rule may be intuitively appealing, but its fatal flaw is in ignoring the circumstances motivating the trades. If an insider's sale has positive implications for corporate control, that sale can actually be *good* news for the firm. This is the sort of sale we consider in this paper.

Of course, some insider sales are bad news. Seyhun (1986), for example, finds that insider sales in his sample are followed by significant declines in their firms' stock prices. He asserts that insiders not only know when the market has mispriced their firms' stock, but that they take advantage of that mispricing. Lorie and Niederhoffer (1968), Pratt and DeVere (1970), Jaffe (1974), and Finnerty (1976) also present evidence suggesting that insider trades generate significant abnormal profits. Such evidence bolsters the negative interpretation of insider sales.

The problem with this general interpretation is that there is an important class of insider sales that is associated with share price *increases*: involuntary sales that "emancipate" a firm from the voting domination of a controlling insider. For example, Johnson *et al.* (1985) find significant abnormal stock price increases after the deaths of senior managers whose control had been protected by their founder status and/or their large shareholdings. Similarly, Slovin and Sushka (1993) find significantly positive share price responses to deaths of executives owning more than 10% of their firms' stock—enough stock to entrench these executives, in the authors' view—and that this effect gets stronger, the more stock the insiders held. Demsetz and Lehn (1985) report that the stock prices of Disney, Gulf + Western, and Chock Full O'Nuts rose 25%, 42%, and 22%, respectively, when their "dominant" owners died; Holderness and Sheehan (1988) note that James Crosby's death caused the stock of his "personal fiefdom," Resorts International, to rise from \$49 to \$67.25. Clearly, the market did not interpret these insiders' divestitures as attempts to parlay superior information into trading profits. Instead, the positive implications for corporate control translated these sales into positive changes in firm value.

In stark contrast to the types of sales studied by Seyhun, the "death" studies are about control. The insiders had been entrenched. If they chose to pursue activities that would increase their private benefits at the expense of outside shareholders' value, they could do so with impunity. If we characterize insider sales along two dimensions—voluntary v. involuntary and control potential v. none—we see that

Seyhun's results apply to the voluntary/no-control sales types. However, the contrary stock price response in the death studies (the involuntary/control sales) show that generalizing Seyhun's results to all insider sales ignores ignore crucial mitigating factors such as *who* is selling and *how much*.[2] In this paper, we consider those factors in a model of a third type of insider sale: a voluntary/control sale.

In this model, a manager's willingness to relinquish voting control, thereby exposing himself to meaningful oversight, can be a positive signal of his value. A manager who protects himself from shareholder scrutiny is perceived to be of low quality (like the managers whose deaths emancipate their firms). In the signaling equilibrium presented, the more stock a manager sells beyond a control threshold, the more valuable his shareholders expect him to be.

One of the factors influencing the manager's sale choice is the makeup of the shareholder base that would be newly able to monitor her. *To whom* would she be vulnerable? The possible reactions of other blockholders would be particularly important considerations. Most studies of blockholder behavior assume that firms have a single blockholder among a sea of atomistic outsiders. However, given the prevalence of block ownership in American corporations and the evidence that minority blocks as small as 5% can confer significant control,[3] it is likely that many companies have multiple blockholders and that interactions among them can affect corporate control. However, as Holderness (2003) notes, "[s]tudies infrequently address the stock ownership of outside shareholders who do not serve on the board of directors" (p. 53). In this paper, we make a first pass at incorporating multiple blocks in our description of ownership structure. Explicit consideration of the interaction among blockholders is one of the contributions of this model.

The paper is organized as follows. Section 2 presents the basics of the signaling model of control block sales. Section 3 describes a separating equilibrium in which larger sales signal higher-quality managers. Section 4 discusses the model and provides links between it and supporting literature. Section 5 concludes.

THE MODEL

We model a controlling shareholder's decision to sell enough of his stock to become vulnerable to outside oversight. The controlling manager, E, owns the proportion α_0 of his firm, which is enough to ensure his voting control. (We will therefore assume that $\alpha_0 > .50$.) There are two other types of shareholders: an outside blockholder, L, who owns the proportion α_L, and a set of atomistic outside shareholders who own the balance, $[1 - \alpha_0 - \alpha_L]$. E's action in the game is to decide how much, if any, of his stake he will sell to the atomistic shareholders. (We will call this proportion α.) Once he's chosen, the outside shareholders decide whether to challenge him; if he's successfully challenged, he is fired and replaced. The players' payoffs in the game depend upon the state of nature and the identity of the chosen manager (E or his replacement). The game is summarized in the schematic below.

0 1 time

Nature reveals state S to E; E chooses signal α	outsiders may receive signal ι; may choose to challenge E	manager determined; payoffs realized

E bases his sale decision on the state of nature and on his expectations about outside shareholders' reactions to his choice. We will consider the latter influence in the next section, where we describe our specific signaling equilibrium and show how E's decision is a best response to the market's beliefs. In this section, we will describe the more general aspects of the game.

The state of nature influences E because it determines the value of the firm under his leadership. The state is high, medium, or low ($S \in \{H,M,L\}$), with higher firm values possible in higher states ($V^S \in \{V^H, V^M, V^L\}$; $V^H > V^M > V^L$). However, these higher values depend upon E's leadership: if he is fired, firm value is certain to be only V^L. E is uniquely able to generate value from the firm's assets in higher states. At the model's time 0, he learns the state, becoming perfectly informed about his marginal contribution to value.

In addition to being uniquely able to contribute to the firm's value, E is also uniquely able to extract his resources. We model his state-dependent compensation as b(S), where b(H) > b(M) > b(L) > 0. Should he be fired, his compensation at any other firm, and outside managers' compensation at his, would be zero. b(S) is meant to represent all elements of E's compensation. For example, b(S) incorporates Shleifer and Vishny's (1986) definition of compensation as "all transfers from shareholders that the manager negotiates with the board, including direct monetary compensation, expenditures on perquisites such as airplanes and charity, and pet projects the board accedes to while knowing they are wasteful" (p. 128).[4] These sorts of benefits could accrue to a controlling manager as a consequence of his voting control, for example, or from wage contract required by the founder when the firm was initially taken public.[5] However, b(S) may also include increased "leniency and lack of oversight by the board,"[6] especially if outside shareholders perceive that E's expertise is contributing to value. This sort of leniency can be valuable to E, even when he is majority holder, since shareholders could still affect his access to resources through different intensities of monitoring. (The activities of H. Ross Perot, Kirk Kerkorian, and Carl Icahn are examples of the potential influence outside shareholders can exert.) Wherever they come from, the benefits b(S) are unavailable at any other firm; E therefore must consider the value of this compensation when choosing his share-sale signal.

If E keeps his job and receives benefits, net firm value will be as follows:

$$V^H - b(H) > V^L \quad (1)$$

$$V^M - b(M) = V^L \quad (2)$$

$$0 < V^L - b(L) < V^L \quad (3)$$

In the low state, allowing benefits b(L) means that firm value is lower under E than it would be under his replacement. In the medium state, E is able to capture all of his marginal contribution to value; outsiders are different between his leadership and his replacement's. However, in the high state, some of E's contribution is shared with the outside shareholders (since (1) and (2) imply that $[V^H - V^M] > [b(H) - b(M)]$). Outsiders wish to retain E in this case, which will critically influence their choice of actions in the game.

The outside shareholders have two possible decisions to make in the game: first, they must decide if they should challenge E's leadership; second, they must vote on his ouster if they decide to challenge. Given (1), (2) and (3) above, it is obvious that:

if S=H, outside shareholders would not want to challenge E

if S=M, outside shareholders are indifferent to challenge

if S=L, outside shareholders would want to challenge E.

However, unlike E, outsiders do not know the state. Instead, they must update their priors based on information revealed during the game. Outsiders receive one or two signals. The first of these signals is E's share sale itself. The second is a noisy signal from nature about the state, which we will call ι. Outsiders will only receive this second signal if E sells enough stock to make himself vulnerable; if instead he chooses to retain voting control, no ι signal is provided.

Even if they receive the ι signal, outside shareholders are at a disadvantage relative to E. Either because they lack access to some relevant information, or because they lack the expertise to fully evaluate it, outsiders cannot perfectly distinguish the state of nature. Instead, their ι signal takes on only two values: *h* and *l* (high and low). It is always *h* in the high state and *l* in the low state. However, it can take on either value in the medium state; in this case, $\iota=h$ with probability h^M and $\iota=l$ with probability $l^M = [1 - h^M]$. (That is, ι is state-dependent: the probability, *p*, that $\iota=h$, given that S=H, is 1; however, $p(\iota=h|S=L) = 0$, and $p(\iota=h|S=M) = h^M$.) Thus, outsiders cannot distinguish between the medium and high states, given $\iota=h$, or between the medium and low states, given $\iota=l$. However, given their incentives, it is clear that the outsiders' best strategy is to:

challenge E if $\iota=l$ (since S = L or M)

do not challenge E if $\iota=h$. (since S = H or M).

This clarifies outsiders' evaluation of the first of their two decisions in the game.

Their second possible decision is the firing decision. If they challenge E, outsiders must then vote to keep or fire him. In our simple approach to modeling the interactions among the blockholders, we assume that the outside blockholder, L, is hostile and votes all of his α_L shares against E. E will then be fired if enough of the atomistic shareholders also vote against him. We assume that a given small shareholder is more likely to vote to fire in lower states of nature. Following Stulz (1988), we use the proportion s(S, α) to describe this voting behavior; s is distributed uniformly between d(S) and 1, and is larger in lower states (d(L) > d(M) > d(H)). Given this voting behavior, E will be fired if:

$$\alpha_L + s(S, \alpha)*(1 - \alpha_0 + \alpha - \alpha_L) > .50;$$

that is, if

$$s(S, \alpha) > (.5 - \alpha_L)/(1 - \alpha_0 + \alpha - \alpha_L) \equiv z(\alpha).$$

At least the proportion z(α) of the atomistic shareholders must vote against E for him to be fired. The probability of a successful challenge is therefore the probability that s(S, α) exceeds this minimum:

$$p[\text{fire}|\text{challenge}] = p[s(S, \alpha) > z(\alpha)] = [1 - z(\alpha)]/[1 - d(S)] \equiv F(S, \alpha).$$

Thus, E is more vulnerable the smaller is his initial block (α_0) and the larger is the block of the hostile outsider (α_L).

Outsiders' two decisions in the game stem from their incentive to try to get rid of E if they think they would be better off with another manager. E must consider this incentive when determining his own action in the game, the amount of stock he will sell. He can only be challenged by his shareholders if he gives them the opportunity—that is, if he sells enough stock. He therefore will only risk a challenge if taking that risk makes him better off.

E makes his α choice after he learns the state of nature, and his choice maximizes his expected wealth, given that state. His expected wealth depends on three things: the value of his post-sale holdings in the firm (the proportion $[\alpha_0 - \alpha]$); the proceeds from any share sales; amount of his compensation, b. His objective function takes the following form:

$$\max E_0^{E}(W_1|S) = (\alpha_0 - \alpha)*E_0^{E}\{V^S - E_0^{M}[b|\iota, \alpha]|S, \alpha\} + \alpha*E_0^{E}\{E_0^{M}(V^S - b|\iota, \alpha)|S, \alpha\} + E_0^{E}\{E_0^{M}(b|\iota, \alpha)|S, \alpha\}. \quad (4)$$

(The *M* superscript on a variable indicates that the argument depends on shareholders' perception of the state, which is not necessarily the true state.) We can clarify the tensions driving E's actions by rearranging (4) this way:

$$E_0^{E}(W_1|S) = \alpha_0*E_0^{E}(V^S|S, \alpha) + \alpha*\{E_0^{E}[E_0^{M}(V^S|\iota, \alpha)|S, \alpha] - E_0^{E}(V^S|S, \alpha)\} + (1- \alpha_0)*E_0^{E}\{E_0^{M}(b|\iota, \alpha)|S, \alpha)\}. \quad (5)$$

The first term of equation (5) represents the value of E's shares. This depends both on the state and on the manager: if E keeps his job, his firm will be more valuable in higher states, but if he is fired, it will only be worth V^L. If he chooses to become vulnerable, he may lose his job and sacrifice his positive marginal contribution to value.

The second term in equation (5) represents E's trading profits. As we will see below, as long as he takes actions along the equilibrium path—signals truthfully—these profits will be zero. However, he may be tempted to falsely signal a higher state, gambling that he will keep his job, generate trading profits, and receive higher benefits. In order for the signaling equilibrium to obtain, any expected gains from such a false signal must be outweighed by the expected costs of losing his job.

Those costs include losing all of his benefits, b(S). The third term in equation (5) represents these benefits (adjusted for E's own contribution to them as a shareholder himself). In order to receive any benefits, E must convince his outside shareholders that he is more valuable than any potential replacement—that is, that the state is not low. Signaling a higher state, however, means becoming vulnerable. Again, this is the primary tension driving the model: in order to increase his benefits, E must risk losing his job, which would eliminate all of his own marginal compensation and doom his firm to its lowest possible (gross) value.

This section has described the basics of the signaling game played by E and his shareholders. We can summarize this game as follows. At time 0, Nature reveals unambiguously to E what time 1 firm value will be under his leadership (V^S); E must then decide what proportion of his shares to sell (α). He will choose the α that maximizes his expected wealth, considering his share ownership, his managerial compensation, and his trading profits. Shareholders then use E's action, along with any ι signal from Nature, to update their priors over the states and to decide whether to challenge E's leadership. If they successfully challenge him, time 1 firm value will be V^L, and managerial compensation will be zero. In all other cases, time 1 value is V^S and managerial compensation is positive; however, E's marginal contribution to shareholder wealth can be positive, negative, or zero, depending on the state.

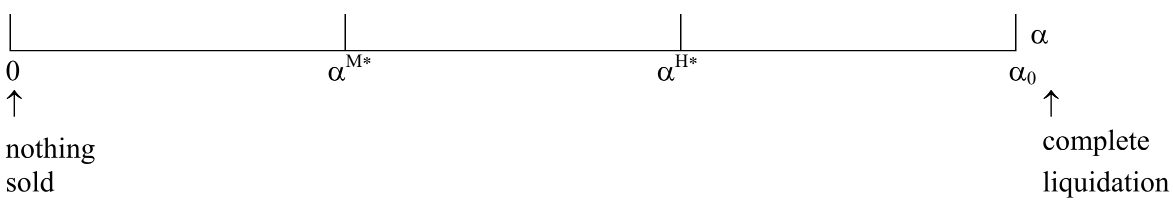
Having described the basics of the model, we now go on to consider in detail a potential separating signaling equilibrium, in which E signals higher states with higher share sales.

A SEPARATING SIGNALING EQUILIBRIUM

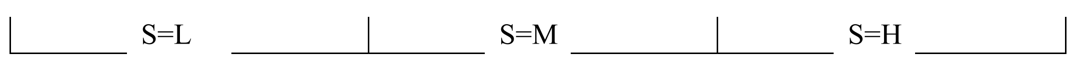
Model Description and Development

To establish an equilibrium, we must specify a self-sustaining set of actions in which both players' choices are a best response to the action of the other player. To describe such a set, we will first specify a set of beliefs that govern the choices of the outside shareholders. These beliefs must be sustainable in the sense that they are rational, given the outsiders' information set. We then demonstrate that E's best responses to the shareholders' actions cause their beliefs to be self-fulfilling; the chosen strategy for each player is then optimal given the strategy of the other, and the equilibrium is established.

In the signaling equilibrium we consider, E sells more shares in higher states. Outsiders believe that a higher-valued manager does not need to protect his job with voting control; only a low-valued manager would be afraid of scrutiny. Outsiders codify their beliefs by translating E's share sales as follows:



market's beliefs:



α is in range described as:

Thus, if E sells an amount less than α^{M*}, shareholders believe that S=L and will allow E compensation of only b(L) as long as he is manager; on the other hand, if E signals by selling an amount greater than α^{H*}, he will receive b(H) if he keeps his job.

In the proposed equilibrium, outsiders set the α^{M*} and α^{H*} bounds so that if E inconsistently signals a state higher than the true state, he will be fired *if he is challenged*. (A schematic illustrating our proposed equilibrium is presented in Figure 1.) For example, using these bounds, falsely signaling the high state ensures that a challenged medium-state manager will be fired (that is, α^{H*} sets $F(\alpha^{H*}, M) = 1$). We can solve for this signal by setting the minimum proportion of outsider votes against E in the medium state, d(M), equal to the proportion required for ouster, $z(\alpha^{H*})$:[7]

$$z(\alpha^{H*}) \quad = \quad (.5 - \alpha_L)/(1 - \alpha_0 + \alpha^{H*} - \alpha_L) \quad \equiv \quad d(M).$$

This equality implies that:

$$[(.5 - \alpha_L) - d(M)*(1 - \alpha_0 - \alpha_L)]/d(M) \quad = \quad \alpha^{H*}.$$ [8]

Similarly, to ensure that a low-valued manager will be fired if he chooses a medium-state signal, we set $z(\alpha^{M*}) = d(L)$, which implies that:

$$[(.5 - \alpha_L) - d(L)*(1 - \alpha_0 - \alpha_L)]/d(L) \quad = \quad \alpha^{M*}.$$

Because $d(L) > d(M)$, α^{H*} is always greater than α^{M*}, so that E must sell more stock if he wishes to signal the higher state. Also, since $(\alpha_0 - \alpha^{M*}) < .50$, signaling either the high or medium state forces E to relinquish majority ownership.[9] If this were not so, there would be no risk to falsely signaling a higher state—no cost to a truthful signal—since the signal would not leave E vulnerable.

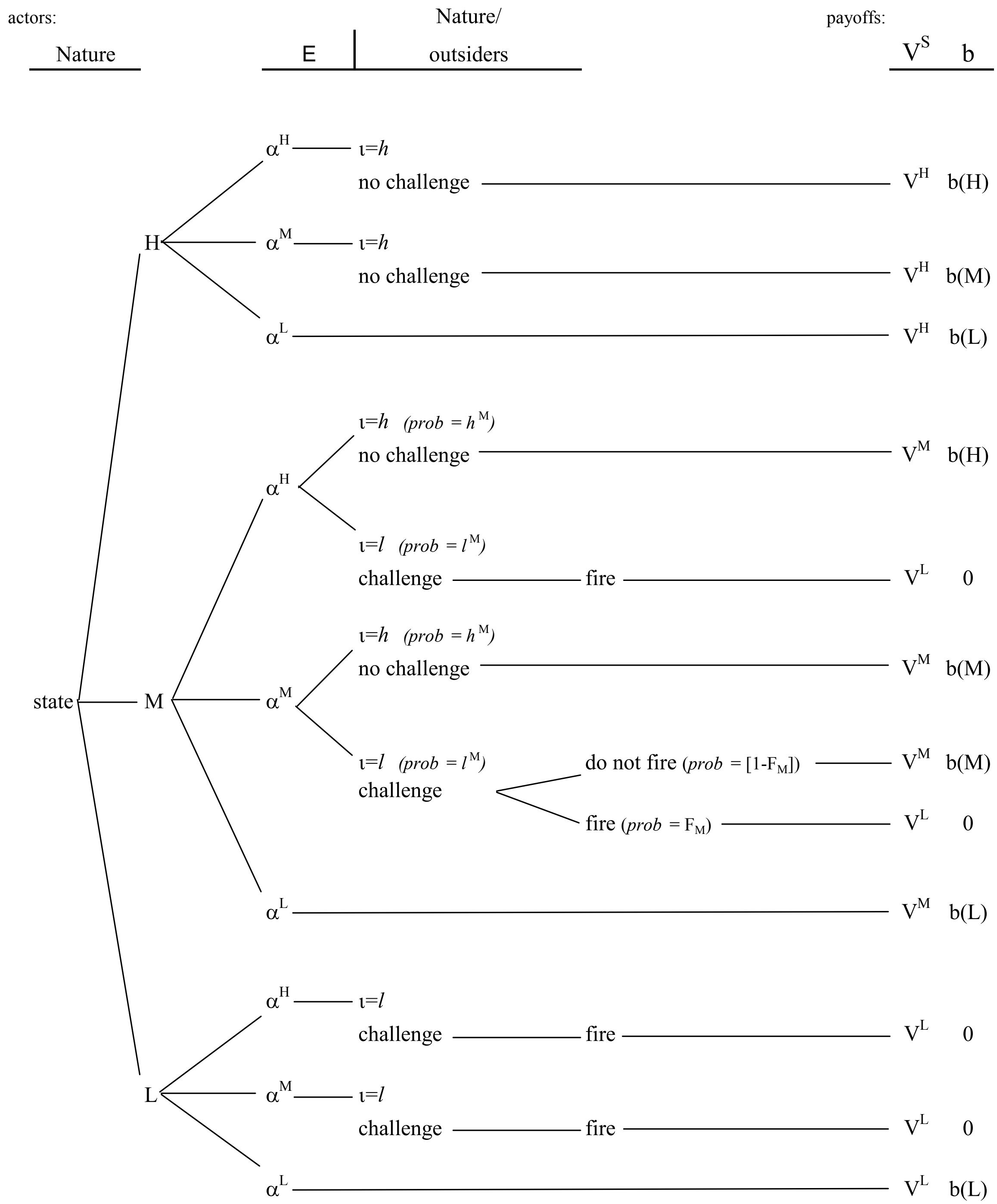
To finish our description of the market's beliefs, we must specify their interpretation of out-of-equilibrium actions by E. Some of these actions are easily detected by outsiders, since they must be inconsistent with outsiders' exogenous ι signal. For example, if a low-valued manager signals the high state, outsiders will receive an inconsistent $\iota=l$ signal; outsiders will then know that the true state is either medium or low. Similarly, if a high-valued manager signals the low state, their inconsistent $\iota=h$ signal will tell the outsiders that the state is actually medium or high. In order for their beliefs to be sustainable in these cases, outsiders' updating must consider only the states that are consistent with their observation of ι. Consistent with Welch (1989), we will specify that outsiders assume the worst when E sends an inconsistent signal: they assume that S=M when $\iota=h$, and that S=L when $\iota=l$.

Having described the market's beliefs, we must now show that E maximizes his expected wealth when his actions are consistent with those beliefs. We will then have established the separating signaling equilibrium. Figure 1 helps us visualize the necessary comparisons.

For both the low and high states, a consistent signal clearly dominates E's choices. A low-state manager has only one way to receive positive compensation: keeping his job. However, inconsistently signaling that S=M or S=H means getting fired. Only by choosing the consistent α^L signal will he earn b(L) and maximize his expected wealth. On the other hand, in the high state, E knows he will never be challenged, since outsiders are certain to receive the exogenous signal $\iota=h$. Thus, if he were to choose not to signal S=H, he would simply lower both his expected compensation (b(H)) and his trading price on every share he sells. Again, he maximizes his expected wealth by choosing the consistent signal, α^H.

E's choice is not so clear in the medium state. When S=M, a consistent signal makes E vulnerable. However, unlike in the S=H case, this vulnerability actually means something: only the proportion $z(\alpha^{M}*)$ of outsiders must vote against him for him to be fired, and $z(\alpha^{M}*) \equiv d(L) < 1$. Thus, E risks losing his job if he signals consistently. (To simplify the exposition below, we define this probability that a medium-valued manager will be fired if he signals consistently as F_M [so that $F(\alpha^{M*}, M) \equiv F_M$]; substituting, we find that F_M simplifies to $[1-d(L)]/[1-d(M)]$.)

Figure 1: Schematic Tree Illustrating Separating Signaling Equilibrium



To complete the demonstration of the proposed equilibrium, we must show that choosing $\alpha = \alpha^{M*}$ is E's best response when S=M, despite this risk. Table 1 below gives the value of his objective function (equation (5)) for each of his three possible actions. To help clarify the relevant trade-offs, we will consider each of these choices in turn.

Table 1: E's Expected Wealth, Given S=M ($E_0^E(W_1|M, \alpha)$)

α Signal	Expected Wealth if $\iota = l$	Expected Wealth if $\iota = h$
$\alpha = \alpha^H$	$\alpha_0 * V^L$	$\alpha_0 * V^M + \alpha^H * (V^H - V^M) + (1 - \alpha_0) * b(H)$
$\alpha = \alpha^M$	$\alpha_0 * V^M + (1 - \alpha_0 - F_M) * b(M)$	$\alpha_0 * V^M + (1 - \alpha_0) * b(M)$
$\alpha = \alpha^L$	$\alpha_0 * V^M + (1 - \alpha_0) * b(L)$	$\alpha_0 * V^M + (1 - \alpha_0) * b(L)$

First, consider a medium-state manager who chooses to signal the low state, selling α^L. This would force him to take a loss on every share he sells, since shareholders will set their price given both E's α^L signal and their own received ι. If $\iota = l$, E's signal is confirmed, and shareholders expect that $V^S = V^L$. If, however, they receive $\iota = h$, they expect the lower value consistent with h, V^M. E's expected price is therefore between V^L and V^M. However, since outsiders assume that the probability that $V^S = V^L$ is 1 if $\iota = l$, while E knows that value will be low only if $\iota = l$ *and* he is actually fired, outsiders will determine a lower price than is warranted.

In addition to these trading losses, the inconsistent α^L signal would also restrict E's benefits to b(L) if he kept his job; if he lost it, of course, he would receive nothing. Thus, since his trading profits are negative for any positive α^L, and since his expected compensation falls with α^L, the optimal level of α^L is zero. Choosing this inconsistent signal then leaves E majority holder, with an expected wealth of:

$$E_0^E(W_1|S=M, \alpha=\alpha^L) = \alpha_0 * (V^M) + (1 - \alpha_0) * b(L). \tag{6}$$

This must be lower than what E would expect from a truthful α^M signal, if the equilibrium is to obtain.

An out-of-equilibrium α^L signal is defensive, since E can be sure to keep his job, to receive positive compensation, and to have shares worth V^M. His other inconsistent action, however, is aggressive: signaling the high state in search of trading profits and excess compensation. This strategy is riskier, though, since signaling α^H means he will be fired—losing all benefits and making his shares worth only V^L—if the market's exogenous information refutes him (if $\iota = l$). Using the payoffs in Table 1, we can see that these trade-offs result in an expected wealth from an inconsistent α^H signal of:

$$E_0^E(W_1|S=M, \alpha=\alpha^H) = h^M * \{\alpha_0 * V^M + \alpha^H * [V^H - V^M] + (1 - \alpha_0) * b(H)\} + l^M * (\alpha_0 * V^L) \tag{7}$$

Again, for our equilibrium, this must be lower than what E expects from a consistent signal.

What would E expect from an α^M signal? If he keeps his job, the firm will be worth V^M and he will receive benefits of b(M); if he is challenged (as he will be if $\iota = l$) *and* fired, firm value is V^L and benefits are zero. Substituting into equation (5), we see that this implies that E's expected wealth is:

$$\alpha_0 * \{V^M - (V^M - V^L) * F(\alpha^M, M) * l^M\} + (1-\alpha_0) * b(M)\{1 - F(\alpha^M, M) * l^M\},$$

which, using (2), simplifies to:

$$E_0^E(W_1|S{=}M, \alpha{=}\alpha^M) \quad = \quad \alpha_0 * (V^M) + b(M) * [1 - \alpha_0 - F(\alpha^M, M) * l^M]. \tag{8}$$

This will be maximized when F is minimized, so if E chooses to signal that S=M, he will do so by selling as few shares as possible (by setting $\alpha = \alpha^{M*}$). However, since $[\alpha_0 - \alpha^{M*}] < .5$, even this minimum sales amount will still require him to relinquish majority ownership and risk being fired.

Having described E's incentives when S=M, we can determine the parameter restrictions that will permit the signaling equilibrium. Since E will prefer a consistent α^M signal to the low signal when (8) > (6), we have the following restriction (after rearranging and utilizing the definition $h^M = [1 - l^M]$):

$$h^M \quad > \quad 1 - [(1 - \alpha_0)/F_M] * [1 - b(L)/b(M)]. \tag{9}$$

Similarly, he will choose α^M over a high signal if (8) > (7), which reduces to the following requirement:

$$h^M \quad < \quad b(M) * (1 - F_M)/\{\alpha_0 * [(V^H - V^M) - (b(H) - b(M))] + b(H) - b(M) * F_M\}. \tag{10}$$

Together, these restrictions characterize the parameter values that permit the equilibrium, and give us the following theorem: A separating signaling equilibrium, in which higher share sales signal states of nature, will exist in this signaling game as long as h^M falls between the bounds described by equations (9) and (10).

DISCUSSION

In this section, we discuss how literature on blockholdings, trading by insiders, and the dynamics of family firms can be related to the comparative statics of the theorem just presented. We focus on the model's implications for state-dependent firm value and benefits, blockholder interactions, and the size of the controlling manager's stake.

Firm Value and Private Benefits

Some of the inequalities in the theorem affirm the obvious: E is more likely to truthfully signal the medium state the lower are the benefits from signaling the low state and the temptations to signal the high state. Thus, the signaling equilibrium is more likely to obtain when b(L), b(H), and V^H are relatively low (so that b(M) and V^M are relatively high).[10] The key motivator here is the benefits, so we will focus on them in this section. However, since benefits are necessarily bounded by firm value, we first briefly note some evidence relating firm value and the willingness of controlling insiders to make significant sales.

Truthful signaling implies that higher-state managers should be more willing than low-valued managers to sell significant amounts of stock. There is some empirical support for this proposition. For example, Demsetz and Villalonga (2001) propose that some of their empirical results may suggest that "management choose[s] to hold fewer shares when firms seem to be doing well" (p. 228). In a specific test of this relationship, Livingston (2002) relates a firm's operating cash flow (a proxy for firm value) to its controlling manager's willingness to sell enough stock to fall below a control threshold (where this threshold is set at three different levels, 5%, 10%, and 25%). Operating income is measured both at its level, as its percentage of total assets, and as its percentage of annual sales. She finds that significant sales are associated with higher operating cash flow.[11] Although the results are not significant at conventional levels, the fact that 88.9% of the regression coefficients are positive is suggestive: good performance alone may help insulate managers from shareholder discipline, making large shareholdings unnecessary. This is especially interesting given Barclay and Holderness's (1989) evidence suggesting that blocks are more valuable in firms with higher cash flow. In this sample, managers appear to be releasing control just when that control would be most valuable.

Higher firm value may imply higher potential control benefits. Our model's defensive behavior, in particular, is consistent with empirical and anecdotal observations that link benefits with control. We now briefly revisit the forms that these benefits may take and some research that relates those benefits to control.

There is a large literature describing control benefits, some of which served as our motivation for the b(S) construct. For example, Demsetz and Lehn (1985) suggest that a controlling manager may derive nonpecuniary income from the "ability to deploy resources to suit [his] personal preferences"; he may also enjoy the "amenity potential" of his current job.[12] He may derive utility from his ability to "exercise authority, dictate strategy, and choose which investments the firm will undertake" (Schulze *et al.*, [2003]). More in keeping with our model, however, are the pecuniary benefits of control—perhaps the ability to "pay himself an excessive salary, negotiate sweetheart deals with other firms he controls, invest in negative net-present-value projects, or simply withdraw corporate funds" (Holderness and Sheehan [1988]). For example, Holderness and Sheehan (1988) find some evidence that individual majority owners who are also CEOs tend to pay themselves higher salaries than do CEOs of non-majority owned firms, even though the majority-owned firms tend to underperform. This ability to extract resources from a firm increases the relative compensation that the manager receives, making his employment there more attractive to him (increasing b(L)). The desire to retain the access to those resources can lead to defensive behavior, consistent with the "death" studies discussed in the introduction. In those cases, b(L) could be interpreted as the amount by which the firms' value rose after the firms were "emancipated."

High private benefits from control discourage managers from releasing control. However, even if b(L) is relatively small (so that b(M) is relatively large, encouraging a medium-valued manager to signal consistently), the manager still must contend with the possibility of being fired. We can see from the theorem that lower values of F_M make the signaling outcome more likely. An important determinant of F_M is the amount of stock owned by the hostile outside blockholder, α_L: unsurprisingly, the more he owns, the less likely it is that the controlling manager will allow himself to become vulnerable. We will now briefly consider how the literature on blockholdings, including that on minority blocks, may inform our model.

Interactions among Blockholders

Most previous research on blockholdings considers only single blocks, despite the evidence that multiple blocks may not be uncommon. For example, Barclay and Holderness (1989) cite a 1984 Securities and Exchange Commission survey that shows that among NYSE, AMEX, and OTC corporations, approximately 20% have "*at least one* nonofficer who owns more than 10% of the common stock, and approximately 15% have at least one officer" who owns that much (emphasis added). These authors also refer to a 1989 study by Mikkelson and Partch, who found an average voting concentration of 20% among officers and directors in their 240-firm sample. Similarly, Demsetz and Lehn (1985), in a sample of 511 firms (a sample "heavily weighted by Fortune 500 firms, precisely the firms that are supposed to suffer from diffuse ownership structures"), find that the five and twenty largest shareholders own an average of 24.8% and 37.7%, respectively, of their firms' stock. It therefore may be quite common to find several blockholders in a single firm.

If multiple blocks are common, then interactions among their owners may be important determinants of firm value. Pagano and Roell (1999) recognize this when they suggest that a controlling blockholder's incentives can be effectively monitored by other large blockholders. Barclay and Holderness (1989) note that "a blockholder's effective control of a corporation will almost certainly be less if he is one of two large-block shareholders than if he is sole blockholder." This control effect may influence a manager's behavior. In Barclay, Holderness, and Pontiff (1993), for example, the authors consider how controlling managers of closed-end fund behave in the face of share concentrations in hostile or friendly hands. The defensive actions taken by the controlling managers in these funds are consistent with those that E may take in our model. For example, if the closed-end fund has a discount, opening the fund would eliminate the discount and increase share value; unfortunately for the outside shareholders, though, controlling managers are presumed to prefer keeping the fund closed, protecting their access to fund resources. For both these fund managers and for E, the temptation to choose such value-decreasing actions is exacerbated by the presence of hostile outside blockholders. However, the outsiders may prevent acting on that temptation. In the closed-end funds, outsiders may accumulate blocks, attempting to amass enough power to take the fund public; similarly, in our model, L's voting of his block facilitates firing a low-valued manager, increasing firm value by releasing b(L).

The role of the outside blockholder L in our model is also consistent with some of the conjectures in the literature about the role and valuation impacts of minority blocks. There are at least three points of contact between this literature and our model. First, there is the incorporation of private costs of control (see, for example, Bolton and Von Thadden [1998]). Control benefits are critical to our model, but if E also faces unique control costs, these may affect not only his desire to sell a substantial amount of stock, but also the market's response to his sale. We have not considered such costs explicitly in the model, and they do not enter into E's objective function. However, such costs, if present, may be incorporated in b(L), which distinguishes E's worst-case situation at his own firm from that in alternative employment. Extending our interpretation of b(L) to include these costs should not change the implications of the model.

The second thread from the minority-block literature that may inform our model is the question of how the minority block was accumulated. Again, our model abstracts from this concern, as we take L's block as a given. Had we modeled outside block formation as endogenous, it is almost impossible to imagine that anyone would undertake to accumulate a block, given the severely restricted liquidity that already characterizes the firm, the stranglehold that E has on private benefits from control, and the impossibility of meaningful monitoring. However, there may nonetheless be a link between our model's signals and those implied by minority-block accumulation.

We appeal here to Hertzel and Smith's (1993) evaluation of discounts on private equity placements, in which they attempt to reconcile the observed discounts on these placements—which, at around 30%, can be substantial—with the resultant positive stock price responses. The authors reason that, if a buyer must incur significant due diligence costs in evaluating his purchase, he may require a discount as compensation. However, his willingness to undertake the purchase sends a good signal to the market, leading to the positive market reaction. The outside blockholder is essentially certifying the quality of the firm through his purchase. This certification is similar in spirit to the signal in our model, in that E's share sale provides a meaningful, costly signal of firm value, which simultaneously opens the door to certification by outsiders. For us, though, just the opportunity for certification is enough. This brings us to the third and most important link between our model and the minority-block literature: monitoring.

Outside blockholders beget monitoring potential. As Demsetz and Villalonga (2001) note, "[t]he greater is the degree to which shares are concentrated in the hands of outside shareholders, the more effectively management behavior should be monitored and disciplined" (p.221). With respect to the minority-block literature, if this monitoring is costly, the discounts observed on minority blocks may reflect a compensating reward. For example, while Hertzel and Smith's (1993) main conclusion about private placements relates to the signaling/information effect just discussed, they also find evidence that the purchasers of the placements perform valuable monitoring functions. They assume that this monitoring is most pronounced in sales to individuals. While in our model E sells not to an individual, but rather to the atomistic shareholders, his sale nonetheless increase the relative size of L's block, resulting in the "material increase in ownership concentration" that Hertzel and Smith associate with enhanced monitoring. Note also that this increase in relative concentration does not require that an outside block be created, which, as noted above, can be costly to the accumulator; rather, since L's block already exists, its increased importance is an immediate consequence of E's sale. E's willingness to accept this monitoring is the very basis for our model's signal.

The sorts of relationships among blockholders that we model through E and L can have special significance in family-owned firms. These firms often have concentrated control, multiple large block shareholdings, and significant competing interests; they may therefore provide useful illustrations of the control and valuation effects of blockholder interactions. For example, Barontini and Caprio [2004] suggest that a family may not be able to act "autonomously" when the firm has other large shareholders. Similarly, Villalonga and Amit (2004) find that having non-family blockholders negatively affects a family firm's value. Even when the blockholders are all family members, there can be conflicts. Schulze *et al.* (2003) describe obstructive behavior that can occur in the "sibling partnership" stage of a family firm's lifecycle (for example, when siblings with similar large stock holdings but different preferences for consumption disagree over the deployment of firm resources, sometimes even paralyzing the firm through "hostage taking"). Family firms, then, may be fruitful candidates for observing E/L–type interactions.

Livingston (2007) provides a direct test of the model's application to family firms. Using three control benchmarks (5%, 10%, and 25%), she runs logit regressions in which the dependent variable is a dichotomous indicator that equals 1 if a manager makes a sale that leaves him below a benchmark (and 0 otherwise). In these tests, family holdings represented our model's outside blockholder L. Rather than consider family members' holdings as substitutes for their own votes, managers in firms with second-generation family members defensively maintained their control, as if their family were a hostile bloc. In fact, over the 14-year study period, managers in these firms, in contrast to their nonfamily counterparts, actually increased their stock holdings, solidifying their control.

The Size of the Control Block

Having discussed results touching on benefits and outside blockholders, we turn now to the size of the manager's own block. The comparative statics on α and on α_0, respectively, give us the most empirically interesting implications of the model: that firm value increases when an owner releases control (and—given the beliefs specified in our equilibrium—increases *more*, the more that owner sells [α]), but that he is less likely to do this, the more stock he starts with (α_0). We will now briefly mention previous work that touches on these two implications.

Being less likely to release control means acting more defensively. Empirical findings that large blocks are unlikely to be broken up are consistent with this type of defensive behavior (see, for example, Denis and Denis [1993]). Barclay and Holderness (1992) provide a link between this behavior and shareholdings, finding that the more stock the largest blockholder owns, the less likely he is to break up his block. These sorts of results broadly support the secondt implication above: that the likelihood of a control sale falls as α_0 rises. In a direct test of this implication, Livingston (2007) uses initial shareholdings to explain a controlling manager's willingness to make a significant control sale (again using the 5%, 10%, and 25% benchmarks). Using data from 81 firms with a controlling manager, as well as from its mutually exclusive family-owned and "nonfamily" subsamples, all coefficients on α_0 were negative. Managers with higher initial stakes are less likely to allow their holdings to fall below a control threshold; larger blocks were less likely to be broken up. As noted above, this tendency was particularly marked in family-owned firms, which were also characterized by higher average initial managerial shareholdings, significantly higher average terminal shareholdings, and significantly higher maximum terminal shareholdings. Managers in these family firms increased their holdings to defend their control.

These sorts of changes in concentration bring us to our model's change variable, α. The evidence on share-price increases following executive deaths—the involuntary/control sales discussed in the introduction—suggest that value can increase when control is released.[13] In fact, Slovin and Sushka's (1993) work also finds that these valuation increases are positively related to the controlling executives' shareholdings. This result is the involuntary analogue to our model's α implication. In a direct test of α's relationship to significant sales, Livingston (2002) presents event-study results from control sales from ten public firms. These sales are defined relative to same three thresholds. Abnormal returns are defined using both a market model and a decile model, and event periods are both one- and two-day windows around significant sales events. The results, while not statistically significant, nonetheless were primarily positive: eleven of twelve test statistics were positive, as were 57% of the firm-level prediction errors. A test on the proportion of positive statistics cannot reject the null hypothesis that π=.50. Consistently with our model's signaling story—or at least inconsistently with conventional wisdom—there was no suggestion whatsoever that the market interpreted these sales as bad news.

CONCLUSION

Voters like politicians' lives to be an "open book": candidates are perceived to be trustworthy if they act as if they have nothing to hide.[14] A controlling manager's willingness to undergo scrutiny could send the same positive signal to outside shareholders. Managers who use unassailable voting control to defend their jobs—and thus their privileged access to corporate resources—may be afraid to let other shareholders determine whether or not their managerial skills compensate for their higher cost. In this paper, we present a signaling model in which higher-valued managers are more willing to release voting control. This positive view of one type of insider sale, the voluntary/control type, runs counter to the conventional wisdom that insider sales are bad news. Not all insider sales are bids for trading profits motivated by negative private information. Instead, sales that are also control events must be evaluated as opportunities to benefit both from increased ownership dispersion and from increased productive monitoring.

ENDNOTES

1. The first four quotations here come from the following four sources, respectively: "Bad News Bulls? How Insider Buying May Be Good," by Serena Ng, *Wall Street Journal*, 11/30/06; David Coleman, editor of Vickers Weekly Insider Report, quoted in "Stock Sales by Insiders Reach High" (*Wall Street Journal*, 9/3/97); Praveen Gottipalli, quoted in "Some Stock Funds Beat Rivals by Following Insiders' Trades" (*Wall Street Journal*, 1/27/97); and Jack Pickler of Prudential Securities, quoted in "VF's Chairman, Two Others Sell Company Stock" (*Wall Street Journal*, 1/28/98). The fifth quotation discusses a strategy of the Schwab Analytics Fund, as described in "Some Stock Funds Beat Rivals by Following Insiders' Trades" (*Wall Street Journal*, 1/27/97).

2. One can very occasionally find recognition of this fact in the popular press. For example, in 1993, the Wall Street Journal quoted "many money managers" as predicting that a breakup of the 57% block owned by Dart, Inc.'s founder Herbert Haft would "spark a big rise" in Dart's shares. Indeed one such analyst suggested that the \$83.50 Dart shares would be worth as much as \$170 if the Hart family holdings were broken up. (*Wall Street Journal*, 8/23/93)

3. On the control potential of small blocks, see, for example, Barclay and Holderness [1991] and Morck, Shleifer, and Vishny [1988]; also note that the SEC's reporting threshold for significant ownership is 5%. On the prevalence of blocks in American corporations, see Barclay and Holderness [1989] and Demsetz and Lehn [1985].

4. For a further discussion of the forms that these activities and benefits can take, see, for example, Livingston (1996), Demsetz and Lehn (1985), Jensen and Ruback (1983), Harris and Raviv (1988a), and Holderness and Sheehan (1988).

5. See, for example, Barclay and Holderness (1989) and Johnson, *et al.* (1985).

6. Shleifer and Vishny (1989), p, 129.

7. Setting α^{H*} in this way is consistent with the outsiders' beliefs about S, given out-of-equilibrium behavior by E (discussed later in the text): given an α^{H*} signal and a contradictory *l* signal, outsiders believe that S=L. Thus, they are better off firing E, since firm value is lower under his leadership than it is under a replacement (V^L - b(L) < V^L).

8. Note that for α^{H*} to be less than α_0, we must have (.5 - α_L)/(1 - α_L) < d(M). This parameter restriction is intuitively plausible, since it implies that the larger is α_L, the smaller can be the proportion of atomistic shareholders voting against E in a successful challenge. We will assume that this inequality holds.

9. ($\alpha_0 - \alpha^{M*}$) < .5 → d(L)/[1 + d(L)] < .5, which is true since .5 < d(L) < 1. Note that this is broadly consistent with Bolton and Von Thadden's (1998) model: for them, "when control is the overriding concern, then even a small reduction in block size below [the proportion that ensures control] involves a discrete upward jump in costs of control loss" (p. 18).

10. In Livingston (1996), we describe the low-end pooling equilibrium in which all managers keep voting control.

11. Firms with higher capital expenditures were less likely to have managers who made significant sales. Tests using operating income were mixed.

12. See also Jensen and Ruback (1983), Spence (1973), Barclay, Holderness, and Pontiff (1993), and Harris and Raviv (1988a). Williams and Linder (2002) provide an example of the recognition of the value of control from the professional literature. They assert that "[i]t stands to reason that blocks of stock than cannot control the direction of the company… would be less valuable than stock that does" (p.27); they then go on to suggest that the appropriate discount for small blocks relative to controlling blocks is 23%.

13. There is also evidence that firm value can increase simply as ownership dispersion increases, even if there is not "control event" involved. (See Slovin, Sushka, and Lai [2000].)

14. I thank Larry Schall for this analogy.

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SURROGATE INVESTMENT STRATEGY: THE CASE OF SPAIN FOR LATIN AMERICA

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ABSTRACT

This study analyzes a surrogate investment strategy by using a developed market as a possible candidate for investment in developing markets. It examines the markets of Spain and four Latin American countries: Argentina, Brazil, Mexico and Chile. Both short-run and long-run relationships are analyzed in this paper by using vector autoregression (VAR) and cointegration methodology respectively. It is found that Spain is affected by the Latin American countries in question, but does not affect them. Thus, it has exposure to these markets. This relationship is also maintained in the long run. Thus, Spain serves as an excellent surrogate for investment in Latin America.

INTRODUCTION

Financial diversification is an important tenet in portfolio investments. Recently stocks in developing markets have performed handsomely compared to the US markets. However, it is not always possible for investors to find low-cost diversified investments in these developing markets. Also, there are cases where there are restrictions on investing in these markets. Managed accounts, such as pension funds, have specific restrictions on investing in emerging markets. These restrictions may be for the protection of principal or a fiduciary responsibility that the mangers have to the fund's beneficiaries. In some instances, rather than being voluntary, the restrictions may be due to political pressures or legal requirements. Such restrictions could cause significant financial underperformance. For example, in the case of California's public employees' retirement system, fund managers are restricted based on country and market criteria. Two-thirds of such funds have underperformed their peers when the peers did not have such restraints (Chernoff, 2007). In these cases, managers facing such constraints could use a surrogate investment strategy to get exposure and yet provide security to investments. Thus, it would be beneficial to find a surrogate market that has exposure to these markets.

This paper investigates one such relationship between Spain and four Latin American markets. It hypothesizes that Spain with its cultural and economic ties has exposure to these markets and is affected by these markets. Hence, Spain would act as an excellent alternative to investment in Latin American markets. Past studies have investigated international (both interregional and regional) linkages. However, none of them have investigated such linkages from the perspective of the exposure they provide to a particular market and the resulting diversification into a particular region. Such an investigation might provide many benefits to investors.

First, by investing in a surrogate, the cost of investing is reduced because of the fewer number of investments. Second, if a surrogate is chosen from a developed market, there is a possibility for investors to reduce risks. Third, it allows them access to markets that would have not been available due to possible restriction of investments in these markets as explained earlier. In all, it will create a unique strategy for investors seeking international diversification. Such an approach, if widely accepted, can help the financial industry develop and market products that allow for surrogate investment. The benefits and insights of this study can also be extended to other markets and regions in the world.

In the case of Spain and Latin America, Spain has had political and cultural ties with the region for centuries. Colonization of most of Latin America by Spain has resulted in its cultural influence on the region. The main influence is the Spanish language, which is spoken by most Latin American countries (except Brazil, which speaks Portuguese). According to Chislett (2002), the upper strata of the society in Latin America share a similar lifestyle to that of Spain. Culture and language have been cited as reasons for expanding into Latin America by companies holding Spanish assets (Chislett, 2002).

Spain has also had a political influence on Latin America, as suggested by Toral (2006a). He cites journalistic work regarding political agreements between Spain and Latin American governments that may have led to some of the privatization deals in Latin America. Privatization of state-owned companies is another similarity between Spain and Latin America, with Spain preceding Latin America and thus being able to share its experience with the region (Toral, 2006b). He further mentions that some companies that have Spanish partners or relationships with Spanish banks follow these banks or partners to Latin America. He cites culture as a reason for Spanish investment in Latin America over and above Central and Eastern Europe, Africa, and Asia. The 1990s saw a marked increase in outside investment in Latin America. This was due to the liberalization process in Latin America during that time period. Spanish involvement in the Euro zone allows it to gain leverage while investing in Latin America. On the other hand, many Latin American companies are able to source Euro capital using Spanish capital markets.

Spanish companies, trying to capitalize on the political, cultural, and linguistic edge that they have over other foreign investors in the region, have been heavily investing in the region in the 1990s. Companies such as Santander, Central Hispano, Banco Bilbao Vizcaya Argentaria, Repsol-YPF Telefonica, and Endesa that have invested in the region represent about 70% of the trading on the Madrid stock exchange (Vitzthum, 2003). Hence, any fluctuations that happen in Latin America can be easily captured by these companies and the Spanish stock market. This exposure would provide a basis for Spain being a surrogate for Latin American markets. Looking just at short-term correlations, one may not appreciate the exposure of Spain to these Latin American markets. However, long-run relationships along with short-term effects might help one appreciate this link between these markets. Such an appreciation would help open up a whole new avenue for investments.

Finally, this study indirectly (tangentially) supports and proves the benefits of globalization. Companies, by doing business in various countries, including countries that prohibit portfolio investments, are exposed to the markets of these countries. In turn, investors investing in the home markets of these companies get an indirect investment exposure to restricted markets. Thus, globalization that allows multinational companies (MNCs) to do foreign direct investment (FDI) also provides a diversification benefit to investors, thereby reducing their risks. Results show that the Spanish market is affected by Brazil, Chile and Mexico, indicating exposure to these markets. Also, it has a long-run relationship with these markets. Thus Spain may serve as a good surrogate for investment in Latin American markets.

LITERATURE REVIEW

The idea of diversifying internationally stems from the fact that international markets do not behave in lock step fashion. Thus, it is possible for US investors to reduce risk by investing in foreign markets. Initial work in this area has been provided by Grubel (1968), Levy and Sarnat (1970) and Solnik (1974), among others. Since then many studies have tried to analyze the relationships among various stock markets. Studies such as those by King and Wadhwani (1990), Hamao, Masulis and Ng (1990), Koch and Koch (1991), Chelley-Steeley, Steeley and Pentecost (1998), Richards (1995), and Solnik, Boucrelle and Fur (1996), rather than focusing on a particular region, try to investigate the relationship among countries throughout the world.

There are numerous studies that investigate within region interdependencies around the world. For example Monadjemi and Perry (1996) study the Organisation for Economic Co-operation and Development markets; Wang, Yang and Bessler (2003), Africa; Chelley-Steeley, Steeley and Pentecost

(1998), Europe; Chowdhury (1994), Ng (2002), Dekker, Sen and Young (2001), Ng (2000), Daly (2003) and Treepongkaruna, Gan and AuYong (2003), Asia or Southeast Asia or Pacific Basin; and Bailey, Chan and Chung (2000), Soydemir (2000), Soydemir (2002), Haque, Hassan and Varela (2001), Fernandez-Serrano and Sosvilla-Rivero (2002) and Edwards and Susmel (2003), Ratner, Arbelaez and Leal (1997), Ortiz and Arjona (2001), all study Latin America.

There are studies that explore the relationships between the US and Latin American countries. Fernandez-Serrano and Sosvilla-Rivero (2002) use cointegration to find evidence in favor of a long-run relationship between Brazil, Mexico, and the Dow Jones index before the 1998 turmoil and between Argentinean, Chilean, and Venezuelan indices and the Dow Jones index after 1998. They further suggest that the investor has limited gains from long-term international diversification.

Soydemir (2000) using VAR methodology finds significant links between the US and Mexican stock markets and weaker links between the US, Argentinean, and Brazilian stock markets. This research also demonstrates that these links are consistent with the trade links between the countries and, hence, are more related to economic fundamentals than to irrational contagion effects.

In summary, previous studies have tried to analyze relationships among Latin American markets or their relationship with the US market. This study differs from earlier studies in that it explores the possibility of surrogate investing. Specifically, it examines whether a particular market is exposed to the Latin American country markets based on its cultural and economic ties. By this virtue, an investor in such a market would be exposed to Latin American markets. This would decrease the number of country investments and thus reduce costs. Further, it might open up an alternative strategy to investors trying to expose their portfolios to these markets.

The remainder of this paper is organized as follows: The next section describes the data and methodology. The following two sections deliberate on the empirical results and conclusions.

DATA AND METHODOLOGY

Data for this study includes daily observations of indices compiled by the Morgan Stanley Capital Index (MSCI) for Argentina, Brazil, Chile, Mexico, and Spain from March 15, 1999 to March 14, 2004, totaling 1305 observations. These four Latin American countries have the largest market capitalization at the beginning of the time period of this study and the largest GNP (Chen, Firth, and Rui, 2002). Also, these countries were ranked in the top 30 countries for trade and expansion in 2000 (Sowinski, 2000). The MSCI computes data for developed and emerging markets by including 85% of the free float adjusted market capitalization in each industry group within each country (MSCI online). Although not completely investing in all the shares within a country, the MSCI country index is the best approximation to the market index of a particular country. Daily log returns are obtained by taking logarithms of the indices and then taking first difference of these log prices. Such returns are then used to determine the short-term relationship amongst the indices.

Vector auto regression (VAR) is used to analyze the short-term effects of an individual Latin American country's market on the Spanish market and vice versa. The optimal number of lags is obtained using the Box Ljung statistic and the errors are reduced to white noise. In the case of bidirectional relationships to analyze the effect of the Spanish market on a Latin American market, the Latin American index is treated as a dependent variable, the Spanish index as an independent variable, and all the lags of Spanish index are equated to zero. Rejection of this null hypothesis would imply Spain's effect on that particular Latin American market. The reverse relationship is analyzed with Spain being the dependent variable and the Latin American index the independent variable. The effect of the Latin American market on Spain is analyzed by equating all the lags of the Latin American index to zero. The rejection of this null hypothesis would imply that the Latin American market affects Spain. The equation is as follows:

$$Y_t = \sum_{i=1}^{i=r} Y_{t-i} + \sum_{i=1}^{i=r} X_{t-i} \qquad (1)$$

Where,

Y_t = log returns of dependent variables (Spanish/Latin American index)

X_{t-i} = log returns of independent variables (Latin American/Spanish index)

i = number of lags

The regional effect on any one Latin American market is analyzed by having one Latin American index as the dependent variable and the other Latin American indices, and the Spanish index, as the independent variables. A specific country's effect on a Latin American market is analyzed by having the lags of that index equated to zero. If this null hypothesis is rejected, the Latin American market is affected by the country in question. To analyze the effect of all the countries, including Spain, taken together as a group is analyzed by equating the sum of all the lags of the indices to zero. If the null hypothesis is rejected, all the markets in question affect the Latin American market. Similar relationships are also analyzed with the Spanish index being the dependent variable and only the Latin American indices being the independent variables. The equation is as follows:

$$Y_t = \sum_{i=1}^{i=r} Y_{t-i} + \sum_{n=1,i=1}^{n=k,i=r} X_{n,t-i} \qquad (2)$$

Where,

Y_t = log returns of dependent variables

$X_{n,t-i}$ = log returns of independent variables

i = number of lags

n = number of countries

Further analysis of these relationships involves testing whether such short-term relationships are maintained in the long run. The existence of such long-run relationships is investigated using Johansen (1991) cointegration tests. Investigation of long-run relationships using Johansen's cointegration methodology involves the determination of presence of unit roots (non-stationarity) of variables. The null hypothesis of Dickey Fuller (1981) and Phillips Perrone (1988) tests, which are used to determine non-stationarity, is the presence of unit roots. Rejection of the null hypothesis indicates stationarity of variables. The lag length in Johansen's test is chosen such that errors are reduced to white noise based on the Box Ljung Q statistic for serial correlation in the residuals. The null hypothesis in Johansen's test is that there are at most r cointegrating vectors. When either the λ-max or trace statistic

$$\text{Trace statistic} = -T \sum_{i=r+1}^{p} \ln(1-\lambda_i) \qquad (3)$$

and maximum Eigen value test:

$$\lambda_{\max} = -T \ln(1 - \lambda_{r+1}) \quad (4)$$

is significant, the null hypothesis is rejected in favor of r + 1 cointegrating vectors.

There can be a minimum of zero and a maximum of n (number of variables) cointegrating vectors. Thus in the case of each Latin American country's relationship with Spain (bivariate tests) there can be at most two vectors. In addition, in the case of all Latin American countries and Spain there can be a maximum of five vectors. This is a sequential test starting with the null hypothesis of zero cointegrating vectors. Rejection would indicate one cointegrating vector. This testing is continued with the null hypothesis of an increasingly higher number of cointegrating vectors until the null hypothesis cannot be rejected and thus until no additional cointegrating vectors are found.

RESULTS

Short-Term Relationships

Short-term relationships are investigated using VAR beginning with the bidirectional relationships between the Spanish index and each Latin American country index. First, the effect of the Spanish index on a Latin American country index is investigated by equating the lags of the Spanish index to zero. If this hypothesis is rejected, the Spanish market affects that Latin American market. The reverse effect is investigated by equating the lags of the Latin American index to zero. If this null hypothesis is rejected, the Spanish market is affected by the Latin American market.

The results, as indicated in Table 1, show that Spain affects only Brazil at a 10% level of significance. On the other hand, Spain is affected by all the Latin American countries at a 1% level of significance. Thus, it is evident that the Spanish market incorporates the effects of the Latin American markets.

Table 1: Short Term Bidirectional Relationships between Spain and Each Latin American Country[a]

	Dependent Variables				
Independent Variable[b]	Spain	Argentina	Brazil	Chile	Mexico
Spain	--	1.90	2.54*	0.53	1.94
Argentina	3.53***				
Brazil	12.47***				
Chile	15.01***				
Mexico	18.65***				

** 10% level of significance, ** 5% level of significance, *** 1% level of significance. [a]VAR is used to analyze the short-run relationship between the dependent and independent variable. The optimal number of lags is such that the errors are reduced to white noise based on Box Ljung statistic. [b]The null hypothesis that the dependent variable is not affected by the independent variable is tested by equating all the lags of independent variables to zero. Rejection of the null would imply the independent variable affects the dependent variable individually.*

Further evidence for the effect of the Latin American markets on the Spanish market, individually or as a group, is tested using the multivariate framework. First, one Latin American index is treated as a dependent variable while the other Latin American indices and that of Spain are treated as independent variables. The effect of the independent index on the dependent Latin American index is investigated by equating the lags of the independent index to zero. The group effect is investigated by equating the sum of all the independent variables equal to zero.

The results, as indicated in the Table 2, show that Argentina is affected by Brazil at a 10% level of significance. Brazil is affected by Mexico and Spain at 5 and 10% level of significance respectively. Chile is affected by Brazil at a 1% level of significance. There are no other individual or group effects on the Latin American indices. Spain, on the other, is affected individually by every Latin American index, except Argentina's, at a 1% level of significance. The Latin American indices taken together as a group also affect the Spanish index at a 1% level of significance.

Table 2: Short-Term Relationships for Latin American Countries and Spain[ab]

	F-Values for Dependent Variable				
Independent Variable[c]	Argentina	Brazil	Chile	Mexico	Spain
Argentina		0.29	2.61	0.83	1.19
Brazil	4.25*		5.50***	1.46	5.22***
Chile	3.66	1.46		0.87	4.71***
Mexico	1.78	3.75**	1.66		4.45***
Spain	2.48	2.57*	1.21	1.54	
All except dependent variable	3.07	9.05	2.46	1.31	12.86***

** 10% level of significance, ** 5% level of significance, *** 1% level of significance. aVAR is used to analyze the short-run relationship between the dependent and independent variable. The optimal number of lags is such that the errors are reduced to white noise based on Box Ljung statistic. bThis table investigates the effect that Latin American indices and the Spanish index have on each other individually and as a group. cThe null hypothesis that the dependent variable is not affected by the independent variable is tested by equating all the lags of independent variables to zero. Rejection of the null would imply the independent variable affects the dependent variable individually.*

The above results show that there are few relationships among the Latin American countries. Brazil affects Chile and Argentina, which may be because Brazil is the largest market in the region. As Brazil and Mexico are the two largest markets in the region, they are bound to have an effect on each other. However, since Mexico is also a part of NAFTA, its market may experience other influences rendering the effect of the Brazilian market insignificant. The bidirectional effect between the Spanish and Brazilian markets may be due to the large investment made by Spain in Brazil in the late 1990s (ECLAC, 2000). All the included Latin American markets (except Argentina) affect the Spanish market, implying that all investment by Spanish companies in the region is being reflected in the Spanish market. Thus, Spain serves as an excellent candidate for diversification to an investor who wants exposure to the Latin American markets but is leery of their volatility and hence would like a stable market outside the region.
There is a possibility of the relationships being present in the short-run but disappearing on a long-term basis. In such a case, investors may not get the desired benefit of diversification by being invested in Spain as it is exposed to these markets only in the short run. Hence, it is important to test the validity of the relationships during the long run. Cointegration tests are used for testing long-run relationships.

Long-Term Relationships

Stationarity of variables is investigated using the Dickey Fuller (1981) and Phillips Perrone (1988) tests. Results, as indicated in Table 3, show that the null hypothesis of presence of unit roots cannot be rejected for variables in level, but can be rejected in first differences. Thus, all variables are I(1).

Table 3: Dickey Fuller (DF) and Phillips Perrone (PP) Tests for Unit Roots

	Levels		First Difference	
	DFunit	PPunit	DFunit	PPunit
Argentina	-0.08	-0.03	-36.37***	-36.40***
Brazil	-0.94	-1.08	-31.85***	-31.76***
Chile	-0.53	-0.83	-28.54***	-28.48***
Mexico	-2.47	-2.66	-32.62***	-32.58***
Spain	-1.70	-1.66	-35.63***	-35.66***

***1% level of significance

Long-run relationships among the indices are investigated using Johansen's test. A cointegrating vector is identified when either the trace or λ-max statistic (as described in equations 3 and 4 respectively) is significant. This is a sequential test starting with the null hypothesis of a zero cointegrating vector. The results indicated in Table 4 for all indices analyzed together show that the null hypothesis of zero and one cointegrating vector is rejected at a 10% level of significance. Hence, there are two cointegrating vectors in a system of five indices.

Table 4: Johansen's Cointegration Test Results for Latin American Countries and Spain[abc]

H0 ≤ r	λ-max	Trace
0	39.73*	87.37*
1	31.78*	47.65*
2	10.27	15.87
3	4.02	5.59
4	1.57	1.57

** 10% level of significance [a]Johansen's methodology is used to detect the number of cointegrating vectors. The optimal number of lags are obtained using Box Ljung statistic. Lags are increased until errors are reduced to white noise. [b]A cointegrating vector is recognized when at least one of the two statistics reject the hypothesis of r cointegrating vectors in favor of r+1 cointegrating vectors. [c]This is a sequential test starting with zero cointegrating vectors.*

Table 5: Bivariate Johansen's Cointegration Tests between Each Latin American Country and Spain[abc]

	Argentina			**Brazil**	
H0= r	λ -max	Trace	H0= r	λ –max	Trace
0	10.20	11.33	0	21.24*	23.22*
1	1.13	1.13	1	1.98	1.98
	Chile			**Mexico**	
H0= r	λ -max	Trace	H0= r	λ –max	Trace
0	16.05*	17.31*	0	10.23	14.98*
1	1.26	1.26	1	4.75	4.75

** 10% level of significance. [a]Johansen's methodology is used to detect the number of cointegrating vectors. The optimal number of lags are obtained using Box Ljung statistic. Lags are increased until errors are reduced to white noise. [b]A cointegrating vector is recognized when at least one of the two statistics reject the hypothesis of r cointegrating vectors in favor of r+1 cointegrating vectors. [c]This is a sequential test starting with zero cointegrating vectors.*

These results indicate that Spain retains its exposure to the Latin American markets in the long run and thus are supportive of the short-run results. Such support implies that investors seeking diversification into Latin American markets would be well served using Spain as a surrogate.

CONCLUSION

Investors seek international diversification by investing in foreign assets. The recent performance of emerging markets make their addition to a portfolio desirable. However, high volatility in these markets may cause investors to shy away. An alternative would be for a single market to provide diversification into a region. This paper analyzes whether a single market can be used to obtain such exposure to. For this to happen, the market in question should be affected by the regional markets. Investors would be well served if such an exposure is not just short term, but also long term in nature.

This surrogate investment hypothesis is investigated using the case of Spain and the four Latin American markets of Argentina, Brazil, Chile, and Mexico. Spain made direct investments in these markets when they started liberalizing in the 1990s. Because Spain shares cultural and economic ties with the region, it would serve as an excellent surrogate candidate. Results from this study indicate that the Spanish market is affected by the individual Latin American markets studied (except Argentina's). Thus, an investor seeking diversification into Latin America could do so by investing in Spain. However, an investment may have just short-term exposure. Maintenance of long-run exposure is important, though, and is explored using Johansen's cointegration methodology. Results indicate that short-run relationships are also maintained in the long run.

This indicates that the Spanish market is exposed to and is affected by the Latin American markets. Thus investors seeking exposure to these markets, but concerned about them being from developing countries, can do so by investing in the Spanish market. Such a surrogate investment strategy can also be used in other markets of the world. This is a huge benefit especially to institutional investors who may be restricted from investing in developing country markets. It also provides investors with an option for investing in one market (hence reducing costs) and yet being exposed to multiple markets.

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INTRA-INDUSTRY TRADE BETWEEN THE UNITED STATES AND LATIN AMERICAN COUNTRIES

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ABSTRACT

This paper aims to explain the extent of vertical and horizontal intra-industry trade (IIT) in United State's foreign trade with 20 Latin American countries. It also attempts to identify the country- and industry-specific determinants of vertical and horizontal IIT . One of the main findings is that, with the exception of Mexico, the U.S. trade patterns with rest of Latin American countries are dominated by one-way trade. Another main finding is that the observed increase in intra-industry trade between the United States and Latin America is almost entirely due to two-way trade in vertical differentiation. The third important finding is that, among the country-specific determinants, the level of per capita income and trade intensity are found to affect the shares of all three types of IIT positively while difference in per capita income, difference in economic size, distance, difference in factor endowment, and trade imbalances are found to affect the share of all three types of IIT negatively. Finally, among the industry-specific variables, product differentiation, vertical product differentiation, industry size, and product quality differences are found to have a positive effect while industry concentration is found to have a negative and statistically significant effect on all three types of IIT share.

INTRODUCTION

Since the introduction of the concept of intra-industry trade (IIT) in the 1960s, a large number of theoretical and empirical studies have investigated the determinants of this trade. Intra-industry trade is defined as the simultaneous export and import of commodities of the same industry group. Intra-industry trade describes trade in similar, but slightly differentiated products based on imperfect competition, or trade in close substitutes demanded by consumers in different countries who may have distinct tastes or preferences. As Greenway and Milner (1986) and Greenway and Torstensson (1997) point out, the interest in IIT arose mainly because the traditional theory of comparative costs, dealing with homogenous products, is incapable of explaining the simultaneous exports and imports to a country of the same statistical category. The theoretical studies focused mainly on providing explanations for the existence and development of IIT while empirical studies mainly focused on investigating determinants of IIT, with a small number of studies focusing on IIT aggregation and measurement issues.

The majority of empirical studies have tried to explain the IIT of developed countries due to the availability of detailed trade data for these countries. Some recent studies have also attempted to estimate the extent of horizontal and vertical intra-industry trade as well as identify their determinants. Most of these studies are concentrated on IIT in European countries and only a few are on the U.S. IIT. Some of the previous studies on the U.S. IIT include Clark (2006), Clark and Stanley (2003), Gonzalez and Valez (1993, 1995), Hart and McDonald (1992), and Manrique (1987). Despite the diversity of approaches used by these studies, some consistent results and common features regarding the types of factors influencing IIT have emerged. Studies of bilateral trading arrangements have found that similarity in industrial structure, demand patterns, and size of countries are important country-specific factors while the characteristics of product differentiation and scale economies are important industry-specific factors.

This paper attempts to (a) explain the extent of vertical and horizontal intra-industry trade in the United State's foreign trade with Latin America, and (b) identify the country- and industry-specific determinants of vertical and horizontal intra-industry trade. Trade patterns are identified by breaking up total trade into three trade types: one-way (i.e., inter-industry) trade, two-way (i.e., intra-industry) trade in horizontally differentiated products, and two-way trade in vertically differentiated products. Unlike most other studies on intra-industry trade, this study uses detailed trade data at the 10-digit Harmonized System (HS) industry level and covers a longer and more recent period, 1990-2005.

The remainder of the paper is organized as follows: Section 2 provides a brief discussion of the general performance of international trade of the U.S. with the Latin America during the past sixteen years. Alternative measures of intra-industry trade and the estimated model are discussed in Section 3 while Section 4 presents a discussion of the estimated IIT indices. Section 5 presents and discusses the empirical results of the estimated regression models. Section 6 summarizes the main findings.

GENERAL PERFORMANCE OF U.S. TRADE WITH THE LATIN AMERICA

In this section, we describe the extent, nature and dynamics of trade between the United States and Latin America. Of the 20 trading partners in Latin America, Mexico, Venezuela, and Brazil are the largest trading partners of the United States, accounting for about 14% of total United States merchandise trade with the other 17 Latin American partners accounting for only about 4% of total trade (see Table 1). In 2005, Mexico was the largest Latin American trading partner of the United States, accounting for approximately one eighth of the total merchandise trade of the United States. Brazil and Venezuela are the second and third largest U.S. trading partners in the Western Hemisphere, accounting for about 2.7% of total U.S. merchandise trade. The share of U.S. trade with Latin America increased from 12.4% in 1990 to 17.9% in 2005 (see Table 1). The United States' total trade (exports + imports) with Latin America increased significantly from \$110.1 billion in 1990 to \$461.1 billion in 2005, an annual average increase of about 10.3%. The share of U.S. exports to Latin America, however, increased from 12.5% in 1990 to 20.1% in 2005 while the corresponding share of imports increased marginally from 12.3% to 16.7% during this period (see Table 1).

Of the 20 trading partners in Latin America, 8 countries experienced growth rates of total trade exceeding 10% during the 1990-2005 period. The U.S. trade with Latin America grew at a faster rate relative to its trade with all other countries. However, the U.S. trade with the Latin American trading partners as well as with the rest of the world slowed down significantly during 2000-2005 period, especially after September 11, 2001. It should also be noticed that some of the smaller trading partners, each accounting for less than 1% of the U.S. total merchandise trade, experienced rapid growth rates. United State's international trade with Mexico increased significantly during the 1990-2005 period, especially after the implementation of the NAFTA in 1994. The United States' total trade with Mexico increased significantly from \$58.5 billion in 1990 to \$266.6 billion in 2005, an annual average increase of about 11.6%. Mexican share of U.S. total merchandise trade increased from 6.6% in 1990 to 11.3% in 2005. The share of U.S. exports to Mexico almost doubled during this period, increasing from 7.2% in 1990 to 13.3% in 2005. The share of U.S. imports from Mexico also rose during this period, increasing from 6.1% in 1990 to 10.2% in 2005.

MEASUREMENT OF INTRA-INDUSTRY TRADE

Measures of Intra-Industry Trade

The most widely used measure of intra-industry trade is the Grubel-Lloyd (G-L) index (see Grubel and Lloyd (1975) and Lloyd and Grubel (2003)). While several alternative measures of IIT have been proposed in the literature, perhaps the most widely adopted has been the G-L index. It is considered to be the most appropriate measure for documenting an industry's trade pattern in a single period of time. The G-L index measures the share of IIT of industry i for a given country j as

Table 1: Average Growth and Share of the U.S. Trade with Latin America, 1990-2005
(Average share and annual average growth rate for 1990-2005, %)

Country	Total Trade Share			Exports Share			Imports Share			Average Annual Growth Rate		
	1990	2005	Avg.	1990	2005	Avg,	1990	2005	Avg,	Trade	Exports	Imports
Argentina	0.3	0.3	0.4	0.3	0.5	0.6	0.3	0.3	0.3	9.6	14.3	8.6
Belize	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	6.4	7.3	6.3
Bolivia	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	3.5	4.7	3.9
Brazil	1.5	1.5	1.5	1.3	1.7	1.8	1.6	1.5	1.3	8.2	9.1	8.2
Chile	0.3	0.5	0.4	0.4	0.6	0.5	0.3	0.4	0.3	10.4	9.5	11.9
Colombia	0.6	0.6	0.6	0.5	0.6	0.6	0.6	0.5	0.5	7.6	8.5	7.9
Costa Rica	0.2	0.3	0.3	0.3	0.4	0.3	0.2	0.2	0.3	9.2	9.4	9.4
Dominican Republic	0.4	0.4	0.5	0.4	0.5	0.5	0.4	0.3	0.4	7.2	7.5	7.0
Ecuador	0.2	0.3	0.2	0.2	0.2	0.2	0.3	0.3	0.2	10.2	9.8	11.4
El Salvador	0.1	0.1	0.2	0.1	0.2	0.2	0.0	0.1	0.1	11.6	9.0	15.9
Guatemala	0.2	0.2	0.2	0.2	0.3	0.3	0.2	0.2	0.2	9.6	9.5	9.8
Honduras	0.1	0.3	0.2	0.1	0.4	0.3	0.1	0.2	0.2	13.9	12.8	15.1
Mexico	6.6	11.3	10.0	7.2	13.3	11.1	6.1	10.2	9.2	11.6	10.8	12.5
Nicaragua	0.0	0.1	0.0	0.0	0.1	0.0	0.0	0.1	0.0	25.8	18.6	42.9
Panama	0.1	0.1	0.1	0.2	0.2	0.2	0.0	0.0	0.0	6.1	6.9	2.8
Paraguay	0.0	0.0	0.0	0.1	0.1	0.1	0.0	0.0	0.0	9.0	9.9	3.7
Peru	0.2	0.3	0.2	0.2	0.3	0.2	0.2	0.3	0.2	11.7	8.2	14.5
Suriname	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	6.6	5.7	12.0
Uruguay	0.0	0.0	0.0	0.0	0.0	0.1	0.0	0.0	0.0	10.7	9.6	14.8
Venezuela	1.4	1.6	1.2	0.8	0.7	0.8	1.9	2.0	1.5	9.9	8.5	11.6
Total Latin America	**0.3**	**0.3**	**0.4**	**0.3**	**0.5**	**0.6**	**0.3**	**0.3**	**0.3**	**10.3**	**9.6**	**11.0**
Total All Countries	**100**	**100**	**100**	**100**	**100**	**100**	**100**	**100**	**100**	**7.5**	**5.9**	**8.6**

This table shows the shares of the U.S.-Latin America merchandise trade, exports, and imports and their corresponding rates of growth during 1990-2005. For example, Mexico accounted for 11.3% of the U.S. merchandise trade, 13.3% of the U.S. merchandise exports, and 10.2% of the U.S. merchandise imports in 2005. Source: Authors' calculations based on data from World Trade Atlas Database.

$$IIT_{ij} = 1 - \frac{|X_{ij} - M_{ij}|}{(X_{ij} + M_{ij})} \qquad (1)$$

where X_{ij} and M_{ij} are home country's exports of industry i to country j and home country's imports of industry i from country j, respectively. Thus, IIT_{ij} index in (1) measures the intensity or proportion of intra-industry trade in industry i with country j. If all trade in industry i is intra-industry trade, i.e., $X_{ij} = M_{ij}$, then IIT_{ij} = 1. Similarly, if all trade in industry i is inter-industry trade, i.e., either X_{ij} = 0 or M_{ij} = 0, then IIT_{ij} = 0. Thus, the index of intra-industry trade takes values from 0 to 1 as the extent of intra-industry trade increases, i.e., 0 ≤ IIT_{ij} ≤ 1.

The IIT index in (1) can be modified to measure the intra-industry trade in all products with country j as a weighted measure of the IIT_{ij}'s and can be written as

$$IIT_j = \sum_{i=1}^{n} w_{ij}\left[1 - \frac{|X_{ij} - M_{ij}|}{(X_{ij} + M_{ij})}\right] \quad where \quad w_{ij} = \left[\frac{(X_{ij} + M_{ij})}{\sum_{i=1}^{n}(X_{ij} + M_{ij})}\right] \quad \text{, i.e.,}$$

$$IIT_j = \frac{\sum_{i=1}^{n}(X_{ij} + M_{ij}) - \sum_{i=1}^{n}|X_{ij} - M_{ij}|}{\sum_{i=1}^{n}(X_{ij} + M_{ij})} \qquad (2)$$

where n is the number of industries at a chosen level of aggregation.

Measuring Vertical and Horizontal Intra-Industry Trade

The literature on intra-industry trade increasingly emphasizes the importance of differentiating between horizontal and vertical intra-industry trade. Horizontal intra-industry trade ($HIIT$) is generally defined as the exchange of commodities differentiated by different attributes excluding quality, while vertical intra-industry trade ($VIIT$) is the exchange of commodities characterized by different qualities. This explains why the presence of one or the other has different implications for the trading partners. Horizontal intra-industry trade ($HIIT$) is considered to be of greater relevance to trade among developed countries with high and similar per capita incomes while $VIIT$ is considered to be particularly relevant to trade among unequal trading partners with different income levels. Recent empirical studies, however, show that even among developed countries, vertical IIT are predominant as compared to horizontal IIT (see for example, Greenway et al. (1994) and Athurupane et al. (1999)).

In the evaluation of trade flows, quality analysis is undertaken mainly with the use of unit value indices, which measure the average price of a bundle of items from the same general product grouping. The rationale for using unit value as an indicator of quality is that, assuming perfect information, a variety sold at a higher price must be of higher quality than a variety sold more cheaply. According to Stiglitz (1987), prices will reflect quality even with imperfect information.

In disentangling total IIT into horizontal IIT ($HIIT$) and vertical IIT ($VIIT$), we use unit value information at the 10-digit HS industry level as follows:

$$IIT_i = HIIT_i + VIIT_i \tag{3}$$

where $HIIT_i$ is given by (2) for those products (k) in industry i where unit values of imports (UV_{ki}^m) and exports (UV_{ki}^x) for a particular dispersion factor (α) satisfy the condition,

$$1-\alpha \le \frac{UV_{ki}^x}{UV_{ki}^m} \le 1+\alpha$$

and $VIIT_i$ is given by (2) for those products (k) in industry i where,

$$\frac{UV_{ki}^x}{UV_{ki}^m} < 1-\alpha \qquad \text{or} \qquad \frac{UV_{ki}^x}{UV_{ki}^m} > 1+\alpha$$

where α = 0.15. Typically, trade flows are defined as horizontally differentiated where the spread in the unit value of exports relative to the unit value of imports is less than 15% at the 10-digit HS level. Where relative unit values are outside this range products are considered as vertically differentiated. The presumption is that transport and other freight costs do not cause a difference in export and import unit values by more than this percentage. Although we used three levels of dispersion factor (namely, α = 0.15, 0.20, and 0.25) to calculate the horizontal and vertical IIT, due to the limitation of space we are reporting the results only for α = 0.15. Both Abd-el-Rahman (1991) and Greenaway, Hine and Milner (1994, 1995) demonstrate that increasing the range from 15% to 25% does not radically alter the division of trade into horizontally and vertically differentiated products.

MODEL SPECIFICATION: COUNTRY- AND INDUSTRY-SPECIFIC ANALYSIS

Following Greenway and Milner (1994), Hine, Greenway and Milner (1999), and others, a number of country-specific and industry-specific determinants of the U.S. intra-industry trade are identified as main determinants, drawn from the available theoretical and empirical literature. The determinants identified can be listed as follows:

Country-specific Determinants:

Per Capita Income (PCI): Intra-industry trade with any given trading partner may tend to be higher as per capita income (PCI) of the partner country is higher. According to Greenway and Milner (1994), customer demand at low levels of PCI is generally small and standardized with respect to product characteristics, but with higher PCI, demand will become more complex and differentiated. This will lead to greater demand for differentiated products. On the other hand, if the stage of development can be measured by PCI, a higher PCI then leads to higher intra-industry trade. The effect of this variable, measured as per capita GDP in U.S. dollars on the extent of intra-industry trade, is anticipated to be positive, reflecting enhanced demand for differentiated goods.

Difference in Per Capita Income ($DPCI$): Intra-industry trade will be negatively correlated with differences in per capita income, indicating differences in demand structures and/or differences in resource endowments. If PCI is interpreted as an indicator of demand structure, a greater difference in PCI implies that demand structures have become more dissimilar. This indicates that the potential for intra-industry trade decreases. For trade to exist between two countries, there must in each country be a demand for products of high quality produced by the other. Therefore, when the difference between the per capita incomes of two trading partners is greater, the scope for intra-industry trade tends to be smaller. Following Balassa (1986), Balassa and Bauwens (1987), and Durkin and Krygier (2000), the relative difference in PCI in U.S. dollars, between the U.S. and a given country j, is measured as

$$DPCI_j = 1 + \frac{[w_j \ln w_j + (1 - w_j)\ln(1 - w_j)]}{\ln 2} \tag{4}$$

where $$w_j = \frac{PCI_{US}}{PCI_{US} + PCI_j}$$

Difference in Economic Size ($DGDP$): If the economies of two countries are large, there is more scope for intra-industry trade than in cases where the markets are of very different size. Thus, a greater divergence in economic size between two countries yields a lower volume of intra-industry trade. The relative difference in economic size as measured by GDP, between the U.S. and a given country, is measured in a manner similar to the measurement of difference in per capita income in equation (4).

Distance ($DIST$): Intra-industry trade is negatively correlated with the trade barriers between trading partners, representing the availability and cost of information necessary for trading differentiated products. To account for barriers to trade, this study uses transportation cost. Following Balassa (1986) and Nilsson (1999), since no information is available on transportation cost, the direct-line distance between the U.S. and a given trading partner was used as a proxy.

Difference in Factor Endowment ($DFEND$): Following Martin and Orts (2002), we define the factor endowment differences as,

$$DFEND = \left| \frac{Y_i}{L_i} - \frac{Y_j}{L_j} \right|,$$

where $Y_{i(j)}$ is the level of GDP in country $i\,(j)$ and $L_{i(j)}$ is the total employment of country $i\,(j)$. It can be expected that the smaller the factor endowment difference, the more likely for countries to specialize in horizontally differentiated goods and less likely to specialize in vertically differentiated goods. Thus, we can expect the factor endowment difference to affect horizontal intra-industry trade negatively and vertical intra-industry trade positively.

Trade Orientation (TO): Intra-industry trade will be positively correlated with the country's trade orientation. Following Balassa and Bauwens (1987) and others, TO is defined as the residuals from a regression of per capita trade (PCT) on per capita income (PCI) and population (POP).

$$PCT = (Exports + Imports) / Popolation$$

where exports and imports are measured in millions of U.S. dollars and population is measured in thousands. TO is measured as the residuals from the following regression equation:

$$\ln PCT = \beta_0 + \beta_1 \ln PCI + \beta_2 \ln POP + \varepsilon$$

Trade Intensity ($TINT$): According to Greenway and Milner (1995), the extent of intra-industry trade will be positively correlated with the trade intensity ($TINT$) of the U.S. with a trading partner. As the trade volume with a country increases, there will be more chances for more differentiated products to be traded. $TINT$ is defined as the ratio of the U.S.'s trade volume with a country to its total trade volume.

Trade Imbalance ($TIMB$): Trade imbalance is expected to be negatively correlated with the intra-industry trade. Some recent studies (for example, Lee and Lee (1993), Stone and Lee (1995), and Havrylyshyn and Kuznel (1997)) have also used trade imbalance ($TIMB$) as an additional explanatory variable.

Trade imbalance is measured by

$$TIMB_j = \frac{\left| X_j - M_j \right|}{X_j + M_j},$$

where X_j and M_j are exports and imports of the U.S. to and from country j, and $TIMB_j$ is the measure of trade imbalance with country j.

<u>Industry-Specific Determinants</u>:

Product Differentiation (PD): It is expected that industries with higher degree of product differentiation tend to have higher intra-industry trade shares, as more product variety broadens the basis for intra-industry trade. Following Greenway, Hine and Milner (1994, 1995), we define product differentiation as the number of 10-digit HS industries across 2-digit HS industries for the U.S. trading partners. This measure is expected to affect intra-industry shares positively.

Vertical Product Differentiation (VPD): It is expected that industries with higher degree of vertical product differentiation tend to have higher intra-industry trade shares. Following Clark and Stanley (1999), we use the advertising-to-sales ratio at 2-digit HS industry level to measure vertical product differentiation. This measure is expected to affect intra-industry shares positively.

Industry Concentration ($ICON$): Following Crespo and Fontoura (2005), we use the share of sales of the 4 largest firms in the total sales of the sector as a measure of industry concentration. This is the traditional variable to capture the level of concentration of the market. It can be hypothesized that the possibilities for concentration can be expected to decline with the differentiation of the product. Thus, intra-industry trade will be negatively associated with industry concentration.

Industry Size ($INDSIZE$): The size of the industry is measured as the number of products traded with any given country. It may be presumed that as the number of products traded increases, the volume of trade as well as intra-industry trade will increase. Therefore, we expect a positive coefficient for this variable.

Product Quality Differences ($PRQD$**)**: Following Torstensson (1991), Greenaway, Hine, and Milner (1994), Ballance, Forstner and Sawyer (1992), and Blanes and Martin (2000), we measure product quality differences in product i by the ratio between the unit value of U.S. exports and the unit value of U.S. imports. Product quality is expected to have a positive effect on both horizontal and vertical intra-industry trade.

The estimated model is as follows:

$$\begin{aligned} SIIT_{ij} = \beta_0 + \beta_1 PCI_j + \beta_2 DPCI_j + \beta_3 DGDP_j + \beta_4 DIST_j + \beta_5 DFEND_j + \beta_6 TO_j + \beta_7 TINT_j \\ + \beta_8 TIMB_j + \beta_9 PD_{ij} + \beta_{10} VPD_{ij} + \beta_{11} ICON_{ij} + \beta_{12} INDSIZE_{ij} + \beta_{13} PRQD_{ij} + u_{ij} \end{aligned} \tag{5}$$

where $SIIT_{ij}$ is the share of total IIT in gross trade (exports + imports) of industry i with country j and all the explanatory variables are defined above. We also estimated two other models with the share of horizontal intra-industry trade ($SHIIT_{ij}$) and the share of vertical intra-industry trade ($SVIIT_{ij}$) as the dependent variable. Since these shares take values from 0 to 1, the regression equation may have predicted values for the dependent variable that lie outside the feasible interval. So, to restrict the predicted values between 0 and 1, following Stone and Lee (1995), Caves (1981), Bergstrand (1983), and Loertscher and Wolter (1980), we have used a Logit transformation of the dependent variable. In this case, we estimate the following model:

$$\ln\left[\frac{SIIT_j}{1 - SIIT_j}\right] = \beta Z + u \tag{6}$$

where Z is the vector of explanatory variables including a constant, β is the corresponding vector of coefficients, and u is the random error term.

<u>Data</u>

This study is based on detailed trade data desegregated at 10-digit Harmonized System (HS) industries, covering the period from 1990 to 2005. The 20 countries in Latin America include Argentina, Belize, Bolivia, Brazil, Chile, Colombia, Costa Rica, the Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Nicaragua, Panama, Paraguay, Peru, Suriname, Uruguay, and Venezuela. The trade data was obtained from the Global Trade Information Services (GTIS)'s *World Trade Atlas Database* that uses primary data provided by the U.S. Department of Commerce's Foreign Trade Division.

Data on GDP and PCI are from the International Monetary Fund's *World Economic Outlook Database*. The data on geographic distance ($DIST$) is obtained from the CEPII's distance measures database available at http://www.cepii.fr/anglaisgraph/bdd/distances.htm. Data on industry concentration ($ICON$) is from the *2002 Economic Census*. Data on trade intensity ($TINT$), trade imbalance ($TIMB$), and product quality differences ($PRQD$) are from the Global Trade Information Services (GTIS)'s *World Trade Atlas Database*. Data on vertical product differentiation (VPD), as measured by advertising-to-sales ratio, is from Schonfeld & Assiciates, Inc., *Advertising Ratios and Budgets 2004*. Additional information on trade was taken from the International Monetary Fund's, *Direction of Trade Statistics Yearbook* and U.S. Department of Commerce's International Trade Administration. The data on other relevant variables were taken from the International Monetary Fund's, *International Financial Statistics Yearbook 2005* and the World Bank, *World Development Report 2005*.

ESTIMATION OF INTRA-INDUSTRY TRADE INDICES

In this section, we describe the extent of intra-industry trade between the United States and the Latin American trading partners. A specific problem measuring IIT is the level of desegregation. The scope of IIT and its main components heavily depend on the level of disaggregating. We have estimated the shares of intra-industry trade in United States total trade of detailed products for years 1990-2005, at the 10-digit level of the Harmonized System (HS). The shares of IIT in the U.S. trade with the Latin American trading partners are presented in Table 2.

Table 2: Share of the U.S. Intra-Industry Trade with Latin America, 1990-2005
(Intra-Industry Trade as Percentage of Total Merchandise Trade, %)

Country	1990	1992	1994	1996	1998	2000	2002	2004	2005
Argentina	13.2	14.0	15.8	16.5	19.4	28.8	15.4	14.2	15.0
Belize	0.2	0.6	0.5	3.7	6.6	1.1	5.1	1.2	6.2
Bolivia	0.6	4.0	0.1	0.9	2.1	18.5	1.4	1.9	1.1
Brazil	27.1	19.9	21.5	26.3	27.6	34.2	30.7	29.7	27.7
Chile	6.4	7.4	10.7	14.2	16.3	12.0	10.8	15.1	29.3
Colombia	3.2	5.2	6.7	9.8	7.9	10.0	9.4	9.5	8.4
Costa Rica	10.1	7.6	7.8	9.3	10.1	20.3	35.5	38.1	37.7
Dominican Republic	15.0	14.3	13.6	15.5	14.1	15.7	16.5	20.2	22.3
Ecuador	1.0	2.5	7.4	2.7	8.0	11.1	8.3	4.9	4.4
El Salvador	6.7	4.8	6.6	8.2	8.0	13.1	10.3	7.8	9.5
Guatemala	5.1	3.3	3.7	5.0	3.8	8.7	5.5	3.6	4.4
Honduras	5.3	4.6	5.3	9.5	9.2	11.6	12.5	10.7	15.3
Mexico	35.3	42.5	33.7	43.4	41.0	42.5	42.3	37.8	44.7
Nicaragua	0.0	4.0	0.2	0.7	0.5	0.6	1.0	6.9	9.2
Panama	6.1	4.7	6.3	4.1	6.6	7.8	8.7	12.3	11.5
Paraguay	4.3	0.3	0.4	5.2	1.9	0.3	0.2	8.4	3.3
Peru	5.0	4.0	4.9	7.5	8.5	7.2	6.6	6.1	11.7
Suriname	22.0	20.9	15.9	24.3	29.7	39.8	43.4	39.9	34.6
Uruguay	7.0	7.2	14.2	6.1	4.4	3.2	12.1	8.7	4.3
Venezuela	16.6	13.0	15.8	16.4	33.7	14.1	9.0	7.1	10.9
Total Latin America	**25.6**	**28.7**	**25.1**	**31.8**	**32.6**	**34.8**	**34.3**	**30.2**	**34.6**

This table shows how the share of intra-industry trade has changed between 1990 and 2005.
Source: Authors' calculations based on data from World Trade Atlas Database.

The share of IIT is relatively high only for a handful of countries. Of the 20 countries, only 7 countries had a share exceeding 10% in 1990 and 11 countries had a share exceeding 10% in 2005. This finding is not surprising given the smaller size and the level of development of the majority of these trading partners. Larger trading partners such as Mexico and Brazil have relatively larger share of IIT. Although the *IIT* share increased between 1990 and 2005 for majority of these trading partners, the inter-industry trade continued to be the dominant type of trade. For instance, Mexico's IIT share increased from 35.3% in 1990 to 44.7% in 2005 but the inter-industry share was 55.3% in 2005.

In order to get a full understanding of the level of IIT, it is important to know how common this type of trade is in terms of the number of products traded. The number of products traded and the number of products with IIT are presented in Table 3.

Table 3: Number of Products in U.S. Intra-Industry Trade with Latin America, 1990-2005

	1990			2005		
Country	Total Number of Products Traded	Number of Products with IIT	Percent of Products with IIT	Total Number of Products Traded	Number of Products with IIT	Percent of Products with IIT
Argentina	4,828	399	8.3	6,498	678	10.4
Belize	1,482	4	0.3	1,941	27	1.4
Bolivia	1,313	6	0.5	2,082	36	1.7
Brazil	6,731	1,071	15.9	9,621	1,703	17.7
Chile	4,780	179	3.7	6,183	434	7.0
Colombia	5,630	267	4.7	7,700	613	8.0
Costa Rica	4,455	198	4.4	5,860	464	7.9
Dominican Republic	4,742	228	4.8	6,666	523	7.8
Ecuador	3,326	39	1.2	5,013	283	5.6
El Salvador	3,025	47	1.6	4,610	161	3.5
Guatemala	4,186	82	2.0	5,800	217	3.7
Honduras	3,268	38	1.2	4,843	178	3.7
Mexico	10,566	2,363	22.4	13,825	3,125	22.6
Nicaragua	911	2	0.2	3,081	43	1.4
Panama	4,050	94	2.3	4,753	212	4.5
Paraguay	1,390	6	0.4	1,266	15	1.2
Peru	3,478	73	2.1	5,804	293	5.0
Suriname	1,132	2	0.2	1,802	21	1.2
Uruguay	2,040	34	1.7	2,757	100	3.6
Venezuela	5,809	520	9.0	5,989	433	7.2
Total Latin America	**77,142**	**5,652**	**7.3**	**106,094**	**9,559**	**9.0**

This table shows how the number of products with intra-industry trade has changed between 1990 and 2005. For example, in 1990, Argentina had 399 products with both exports and imports. In 2005, this number increased to 678, indicating an increase of intra-industry trade.
Source: Authors' calculations based on data from World Trade Atlas Database.

The number of products traded varies widely across the Latin American trading partners, as evident in Table 3. Generally, these numbers are larger for larger trading partners, such as Mexico, Brazil, and the Dominican Republic. In 1990, U.S. – Mexico trade activities took place in 10,566 10-digit level industries, of which nearly 22.4% of industries (or 2,363 industries) had some intra-industry trade. By 2005, trade activities increased to some 13,801 10-digit level industries, of which nearly 22.5% of industries (or 3,101 industries) had some intra-industry trade. Although the countries with higher share of IIT tend to have a higher share of products with IIT, product shares are relatively lower than the IIT shares.

The weighted average of the Grubel-Lloyd IIT indices computed using (2) for the years 1990 to 2005, for all Latin American trading partners are presented in Table 4. Although the IIT index in United States' trade with Latin America increased marginally during the period 1990-2005, it is not easy identify any trend for any given country. The IIT indices are not much different when we compare larger trading partners with smaller trading partners. The intensity of intra-industry has remained relatively constant during the period from 1990 to 2005.

Table 4: Grubel-Lloyd Intra-Industry Trade Index for U.S. Trade with Latin America, 1990-2005

Country	1990	1992	1994	1996	1998	2000	2002	2004	2005
Argentina	0.343	0.280	0.266	0.277	0.253	0.259	0.339	0.309	0.294
Belize	0.356	0.430	0.385	0.420	0.247	0.544	0.432	0.283	0.442
Bolivia	0.604	0.316	0.551	0.415	0.264	0.296	0.315	0.356	0.421
Brazil	0.313	0.321	0.313	0.274	0.259	0.279	0.288	0.296	0.312
Chile	0.294	0.287	0.262	0.244	0.225	0.262	0.258	0.283	0.257
Colombia	0.319	0.312	0.283	0.270	0.283	0.299	0.294	0.301	0.281
Costa Rica	0.330	0.304	0.292	0.291	0.295	0.311	0.314	0.299	0.295
Dominican Republic	0.344	0.331	0.307	0.322	0.313	0.316	0.326	0.283	0.303
Ecuador	0.338	0.277	0.318	0.290	0.297	0.308	0.270	0.282	0.305
El Salvador	0.355	0.385	0.363	0.355	0.335	0.315	0.311	0.310	0.298
Guatemala	0.312	0.299	0.315	0.241	0.271	0.304	0.303	0.295	0.295
Honduras	0.335	0.360	0.281	0.248	0.291	0.323	0.303	0.306	0.316
Mexico	0.297	0.269	0.261	0.285	0.281	0.288	0.290	0.290	0.293
Nicaragua	0.567	0.502	0.597	0.366	0.269	0.334	0.291	0.294	0.322
Panama	0.267	0.281	0.262	0.262	0.265	0.297	0.277	0.291	0.284
Paraguay	0.214	0.330	0.408	0.322	0.109	0.336	0.338	0.323	0.311
Peru	0.307	0.356	0.297	0.290	0.250	0.308	0.334	0.337	0.295
Suriname	0.201	0.196	0.276	0.218	0.419	0.428	0.423	0.424	0.471
Uruguay	0.384	0.364	0.321	0.321	0.253	0.353	0.404	0.382	0.342
Venezuela	0.307	0.284	0.302	0.267	0.276	0.245	0.275	0.260	0.236
Total Latin America	**0.339**	**0.324**	**0.333**	**0.299**	**0.273**	**0.320**	**0.319**	**0.310**	**0.319**

This table shows the weighted average of the Grubel-Lloyd IIT indices computed using (2) for the years 1990 to 2005.
Source: Authors' calculations based on data from World Trade Atlas Database.

Having discussed the general trends in IIT, let us now discuss the extent of horizontal and vertical IIT in U.S. – Latin America trade. The shares of horizontal IIT ($HIIT$) and the shares of vertical IIT ($VIIT$) are presented in Table 5. While we used three dispersion factors (α = 15%, α = 20%, and α = 25%) to calculate these shares, due to the limitation of space only the shares for the dispersion factor, α = 15% are presented in these tables. While most other studies use only one dispersion factor, we used three dispersion factors to check the accuracy of estimates.

In the process of calculating these shares, we faced a major obstacle; the unit prices of about 5% of products with IIT were not available making it difficult to identify the product as vertically or horizontally differentiated. As a result, the actual shares of $HIIT$ or $VIIT$ presented in Tables 5 could be slightly underestimated. Despite this limitation, our first finding is that IIT is overwhelmingly vertical (Table 5). The average share of vertical IIT for the entire Latin American region ranged from 70% to 90% during the period 1990-2005. The results also show that the share of vertical IIT is relatively lower for larger trading partners such as Mexico and Brazil. However, most of the total intra-industry trade is vertical. This finding is not surprising; it is consistent with the findings of some recent studies (see, for example, Clark (2006), Clark and Stanley (2003)).

EMPIRICAL RESULTS

We estimate three equations, using as the dependent variable the share of IIT, share of horizontal IIT, and the share of vertical IIT. The models are estimated using country- and industry-specific data for 2004. All the relevant industry-specific variables are measured at the 2-digit HS industry level. Regression results are reported in Table 6. All the variables, with the exception of TO, are expressed in logarithmic form. The first seven independent variables are country-specific variables while the last five independent variables are industry-specific variables.

Table 5: Share of Vertical and Horizontal Intra-Industry Trade with Latin America, 1990-2005
(Vertical and Horizontal Intra-Industry Trade as Percentage of Intra-Industry Trade, %)

	Vertical Intra-Industry Share					Horizontal Intra-Industry Share				
Country	1990	1994	1998	2002	2005	1990	1994	1998	2002	2005
Argentina	92.0	89.2	87.7	84.2	83.3	8.0	10.8	12.3	15.8	16.7
Belize	100.0	88.4	100.0	96.6	96.3	0.0	11.6	0.0	3.4	3.7
Bolivia	100.0	88.1	100.0	70.3	99.2	0.0	11.9	0.0	29.7	0.8
Brazil	65.6	94.0	79.4	87.6	93.1	34.4	6.0	20.6	12.4	6.9
Chile	93.9	96.1	77.9	93.1	84.7	6.1	3.9	22.1	6.9	15.3
Colombia	92.7	69.6	95.4	75.1	83.5	7.3	30.4	4.6	24.9	16.5
Costa Rica	77.3	93.2	95.0	98.6	97.2	22.7	6.8	5.0	1.4	2.8
Dominican Republic	77.5	97.3	87.6	88.4	87.0	22.5	2.7	12.4	11.6	13.0
Ecuador	99.7	90.6	54.6	89.5	97.9	0.3	9.4	45.4	10.5	2.1
El Salvador	99.9	86.3	61.3	66.5	95.9	0.1	13.7	38.7	33.5	4.1
Guatemala	62.2	72.0	93.7	77.4	91.2	37.8	28.0	6.3	22.6	8.8
Honduras	99.6	73.3	60.1	84.3	96.2	0.4	26.7	39.9	15.7	3.8
Mexico	86.2	85.5	78.5	85.3	83.7	13.8	14.5	21.5	14.7	16.3
Nicaragua	98.5	97.5	55.8	100.0	69.6	1.5	2.5	44.2	0.0	30.4
Panama	83.0	54.8	85.6	73.2	91.0	17.0	45.2	14.4	26.8	9.0
Paraguay	99.9	100.0	100.0	100.0	99.8	0.1	0.0	0.0	0.0	0.2
Peru	99.6	90.0	57.7	59.8	96.0	0.4	10.0	42.3	40.2	4.0
Suriname	100.0	100.0	99.9	99.9	99.9	0.0	0.0	0.1	0.1	0.1
Uruguay	38.8	97.9	97.8	96.7	73.2	61.2	2.1	2.2	3.3	26.8
Venezuela	60.5	76.1	92.1	68.7	73.7	39.5	23.9	7.9	31.3	26.3
Total Latin America	**80.5**	**86.1**	**79.3**	**85.2**	**84.5**	**19.5**	**13.9**	**20.7**	**14.8**	**15.5**

These shares are based on a dispersion factor (α) of 15 percent.
Source: Authors' calculations based on data from World Trade Atlas Database.

The results presented in Table 6 confirm the theoretical expectations but some coefficients are not statistically significant. The adjusted R^2 values for the three models are relatively low, ranging from 0.08 to 0.12. However, they are similar to the results of previous studies. Among the country-specific determinants, the level of per capita income is found to affect the shares of all three types of IIT positively but statistically insignificant. The positive coefficient for per capita income indicates that the share of IIT will be higher in trade with high income countries than countries with a lower level of per capita income. These findings are similar to those of earlier empirical studies of total IIT (see, for example, Greenway and Milner, 1995; Clark and Stanley, 2003; Clark, 2006).

Difference in per capita income has a negative effect on all three types of IIT shares; however, none of the coefficients is statistically significant. Similarly, difference in economic size also has a negative effect on all three types of IIT shares but only two are statistically significant. The geographic distance from the U.S. to a given trading partner is also found to have the expected negative effect on intra-industry trade shares. However, it is not statistically significant. This could be due to the relatively close proximity of all trading partners within the Western Hemisphere.

The rest of the country-specific variables, namely, difference in factor endowment, trade orientation, trade intensity, and trade imbalance, also display anticipated signs. However, none of these variables is statistically significant. Among the industry-specific variables, product differentiation is found to have a positive and statistically significant effect on all three types of IIT shares. Similarly, the vertical product differentiation is also found to have a positive effect. Industry concentration is found to have a negative and statistically significant effect on all three types of IIT shares. The industry size has the expected positive effect and is statistically significant. The results for the variable measuring quality differences support the hypothesis that the more differentiated products are in terms of quality, the larger the share of bilateral IIT will be. The coefficient has the expected sign and is statistically significant for total IIT share and vertical IIT share at the 1% level.

Table 6: Determinants of the U.S.-Latin America Intra-Industry Trade
(Heteroskedasticity-corrected *t* -statistics in Parentheses)

Independent Variable	(1) Dependent Variable: *SIIT*	(2) Dependent Variable: *SHIIT*	(3) Dependent Variable: *SVIIT*
Constant	27.591 (0.64)	100.036 (1.66)	152.094 (2.03)
PCI	0.106 (0.32)	0.500 (1.11)	0.103 (0.31)
DPCI	-1.908 (-0.48)	-7.754 (-1.38)	-2.749 (-0.70)
DGDP	-13.903** (-2.36)	-11.332 (-1.27)	-12.491*** (-1.95)
DIST	-0.256 (-1.23)	-0.465 (-1.48)	-0.246 (-1.21)
DFEND	-0.467 (-0.40)	-0.848 (-1.01)	-0.145 (-0.18)
TO	-0.001 (-0.41)	-0.002 (-0.94)	-0.001 (-0.25)
TINT	0.145 (0.15)	0.151 (1.03)	0.011 (0.12)
TIMB	-0.306 (-0.36)	-0.032 (-0.26)	-0.028 (-0.33)
PD	0.272* (5.53)	0.269* (3.63)	0.313* (6.49)
VPD	0.151* (2.87)	0.198** (2.36)	0.109** (2.03)
ICON	-1.222* (-4.37)	-1.278** (-2.33)	-1.073* (-3.75)
INDSIZE	0.272* (5.53)	0.612** (2.02)	0.506* (3.33)
PRQD	0.153* (4.77)	0.047 (0.87)	0.170* (5.38)
Adjusted R^2	0.12	0.08	0.12
n	930	526	890

** significant at the 1% level; ** significant at the 5% level; *** significant at the 10% level.*

Among the industry-specific variables, product differentiation is found to have a positive and statistically significant effect on all three types of IIT shares. Similarly, the vertical product differentiation is also found to have a positive effect. Industry concentration is found to have a negative and statistically significant effect on all three types of IIT shares. The industry size has the expected positive effect and is statistically significant. The results for the variable measuring quality differences support the hypothesis that the more differentiated products are in terms of quality, the larger the share of bilateral IIT will be. The coefficient has the expected sign and is statistically significant for total IIT share and vertical IIT share at the 1% level.

The findings of this study are subject to inevitable limitations. The main difficulty arises from the limitation of data; the industry based statistics are only published at the 2-digit *SIC* (Standard Industry Classification) or NAICS (North American Industry Classification System) levels in the U.S., so this limits the scope of empirical studies. For more reliable results, this exercise should be repeated for different time intervals and the change in the calculated IIT levels should be analyzed. However, despite these considerations, we have identified some important country- and industry-specific determinants of U.S.- Latin America intra-industry trade.

SUMMARY AND CONCLUSIONS

This study analyzes the development of intra-industry and inter-industry trade between the United States and the Latin American countries during the period 1990 to 2005. The main objectives of this paper are to (a) explain the extent of vertical and horizontal intra-industry trade in the United State's foreign trade with the Latin American countries, and (b) identify the country- and industry-specific determinants of vertical and horizontal intra-industry trade. For this purpose, trade patterns are identified by breaking up total trade into three trade types: one-way trade (i.e. inter-industry trade), two-way trade (i.e. intra-industry trade) in horizontally differentiated products, and two-way trade in vertically differentiated products. Unlike most other studies on intra-industry trade, this study uses detailed trade data at the 10-digit Harmonized System (HS) industry level and covers a longer and more recent period, 1990 through 2005. The Grubel-Lloyd intra-industry trade index is used to calculate the intensity of these two types of intra-industry trade.

One of the main finding is that the share of IIT is relatively high only for a handful of countries. Of the 20 countries, only 7 countries had a share exceeding 10% in 1990 and by 2005 only 11 countries had a share exceeding 10%. This finding is not surprising given the smaller size and the level of development of the majority of these trading partners. Larger trading partners such as Mexico and Brazil have relatively larger share of IIT. Although the IIT share increased between 1990 and 2005 for the majority of these trading partners, inter-industry trade continued to be the dominant type of trade.

Another main finding is that the observed increase in intra-industry trade between the U.S. and Latin America is almost entirely due to two-way trade in vertical differentiation. The results also suggest that bilateral trade flows between the United States and Latin America have become more intense indicating that trade relations are strengthening.

Among the country-specific determinants, the level of per capita income and trade intensity are found to affect the shares of all three types of IIT positively, while difference in per capita income, difference in economic size, distance, difference in factor endowment, and trade imbalances are found to affect the share of all three types of IIT negatively.

Among the industry-specific variables, product differentiation, vertical product differentiation, industry size, and product quality differences are found to have a positive effect on all three types of IIT shares.

Industry concentration variable is found to have a negative and statistically significant effect on all three types of IIT share.

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AN EVENT STUDY ANALYSIS OF STOCK PRICE REACTION TO MERGERS OF GREEK INDUSTRIAL AND CONSTRUCTION FIRMS

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ABSTRACT

Using the event study methodology introduced by Brown and Warner (1985) for six Greek industrial and construction firms, we attempt to measure the abnormal returns on stock prices on the day of the acquisition announcement. Estimation period and event period in our market model is -211 -11 -10, +10 respectively. In order to allow for asymmetric effect of news on the abnormal returns we use an E-GARCH model for period -211,-1. Empirical results show that on day t=0, AAR go slightly positive, while CAAR remain positive (0.4% and 1.3% respectively). E-GARCH model results show that good news have a positive effect on abnormal returns, while bad news a marginal negative one.

INTRODUCTION

Since the start of the transition process of the ex-communist countries in Southeast Europe there has been a dramatic increase of FDI (Foreign Direct Investment) in the form of cross border acquisitions. This process, although having started at 1989, is still evolving. Countries like Bulgaria, Romania, Skopje, Serbia/Montenegro, Poland, Hungary, Albania, even Egypt and Jordan, were the recipients of new investments. Many improvements of their economic status took place in the last few years, since these countries need to achieve several strict pre-requisites in order to enter the European Union. With the 2004 E.U. enlargement 10 new countries joined in, namely Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia, Slovenia, Cyprus and Malta, while Bulgaria and Romania entered the E.U. just in the 1/1/2007.

The majority of the FDI come from their neighbouring countries, of which Greece has a leading role, due to the similarities of the economic and political climate that exists in the countries mentioned. Western countries find it rather unpleasant or too risky to invest heavily in the Balkans, since neither geographical distance nor cultural state enables any attempt to do so.

The most important factors that attract FDI from Greece in these countries are low labor costs, the similarities of the bureaucratic system that controls investments and close geographical distance. Cross-border acquisitions have some potential disadvantages, as well. For example, the premium paid for the buy, the kind of information that managers and the market have (insider/outsider information), the expectations of the acquirers and of the market as a whole, which is reflected on the stock's price and the eventual over-evaluation of the acquiring company by the acquirer-company's managers (known as 'the hubris phenomenon'). Mergers and acquisitions continue to emerge strong globally, according to Dealogic's data, a company that thoroughly studies companies' concentrations in any form (be it merger, acquisition, joint venture, conglomerate merger, and so on). Their total value surpassed \$1.1 trillion within the year 2005. This rising trend has commenced about a year and a half ago (in the year 2005). Mergers and acquisitions total value during the year 2005 has risen by \$871 billion comparing with 2004. Some worth-mentioned examples, Guidant, a big company dealing with medical equipment, accepted a bid offer from Boston Scientific, a bid worth of \$27 billion; Mittal Steel's bid over \$18 billion for acquiring Arcelor is still being discussed. In most cases, a company's motive to carry on an acquisition is the search for further development, and the company's optimism about the economy in which it is willing to invest.

In this research, we use the event study methodology in order to determine the effect of the announcement of acquisitions on the average abnormal returns and the abnormal return volatility for Greek construction and industrial firms listed on the Athens Stock Exchange (A.S.E.). Daily data of stock prices is used. According to Brown and Warner (1985), daily data are more accurate than monthly when using the market model. Beyond the methodological issues, the principal results of this study reflect and confirm previous literature.

The rest of this paper is organized as follows: in the 2nd section we briefly review past similar studies on announcement days and on mergers and acquisitions in general, in section 3 we present the sample data in detail and discuss the methodology used, while the 4th section deals with the empirical investigation and results. Finally, 5th section summarises the conclusions.

LITERATURE REVIEW

Sanjin Bhuyan (2002) in his study examines the effect of forward vertical integration on industry profitability by regressing profitability against a number of variables (advertising, value-added per worker, R&D, and so on). Using an input-output methodology, he proves that there exist negative impact profits from mergers. This may be due to the failure of firms to create differential advantages from the acquired firm.

In a study of foreign direct investments towards central and eastern Europe, Balaz Egert, Peter Backe and Tina Zumer (2005) exhibit the level of credit that the private sector of the Balkan economies accept in GDP terms. Generally, the banking sector is the main source of foreign investment. Using a cross sectional analysis and a framework which includes factors driving both the demand for and the supply of private credit, they find that credit growth will very likely remain fast in central and eastern Europe. Moreover, this rapid growth of credit expansion does not pose any risks of deterioration of asset quality.

A. Koulakiotis, N. Papasyriopoulos and Ap. Dasilas (2006), who investigated the effect of the announcement date of acquisitions on the value of stock prices of seven Greek financial firms listed on the Athens Stock Exchange, carried out a similar study. Using the Market adjusted model, GARCH and E-GARCH techniques, they conclude that cumulative abnormal returns start to decline right after the announcement of acquisitions, while the impact of 'bad news' tend to be significant at 15% level of significance.

Annalisa Caruso and Fabrizio Palmucci (2004) use the event study methodology to investigate the market reaction to mergers and acquisitions in the Italian banking sector. They compare the outcomes of using three different dates as the event date, namely rumours, announcement, and outcome date. Interestingly enough, they use 'rumours date' as t=0 instead of the announcement date. Apart from that, they use 4 different models to calculate the ARs: i) the Market return model, $AR_{j,t} = R_{j,t} - R_{M,t}$, ii) the Sector index return model, $AR_{j,t} = R_{j,t} - R_{S,t}$, iii) the Market model expected return with beta calculated with respect to the market index, $AR_{j,t} = R_{j,t} - (\alpha_j + \beta_j R_{M,t})$, and iv) the Market model expected return, with beta calculated with respect to the sector index, $AR_{j,t} = R_{j,t} - (\alpha_j + \beta_j R_{S,t})$. They conclude that, using different event dates will lead any similar observation to different results, while the market believe in the possible value creation from mergers-and-acquisitions operations, but if there is any, it is beneficial to the targets' shareholders and the buyers' management only.

An interesting comparison of event studies methodology and simulation approach was carried out by Thomas Dyckman, Donna Philbrick and Jens Stephan (1984), where they compare 5 different models in order to examine the interaction of portfolio size, event date uncertainty and the magnitude of the abnormal performance from a database of 20,690 observations. Models used were i) the Mean-adjusted returns, ii) the Market-adjusted returns, iii) the Market model using an OLS beta, iv) the Scholes-Williams beta model, which is a variation of the Market model, and v) the Dimson beta model, another variation of the Market model. Parameters for each of the five models were calculated from the period -120,-60 and +60, +120. Event period was -59, +59. Comparison shows that the abilities of the first three models to detect correctly the presence of abnormal performance are similar, with a slight preference for the Market model.

A useful review of the event study methodology since 1969 is available at John J.Binder's paper (1998), where he justifies the reasons why the event study methodology has become the standard method of measuring security price reaction to an announcement or, generally, an event. Event studies are used to test the null hypothesis that the market efficiently incorporates information, while under the maintained hypothesis of market efficiency, they enable the examination of the impact of some event on the wealth of the firm's security holders. Again, the Market model is used. A useful note is that, if there is a great change in the beta coefficient because of the event, AAR will be calculated from a period after that event. Verifying others, he points out that when a large sample of unrelated securities is used or the event dates are not clustered in calendar time, the Market model estimation of the AAR is generally unbiased. Finally, he justifies that non-normality of individual abnormal return and having or not cross-sectional data do not affect the model's performance.

DATA AND METHODOLOGY

We take into account a total of 221 days of stock pricing. The Athens Stock Exchange distributes information through its Daily Price Bulletin and other means details about the prices and the composition of the indexes. Prices of indexes are calculated every 30 seconds during the days of conferencing of the Athens Stock Exchange, using the current stock prices.

Table 1 presents the population of the study, which consists of 10 events of acquisitions of shares that six Greek companies performed. Details about the announcements were taken from the daily press releases of the Athens Stock Exchange. All of the acquiring companies are listed in the Athens Stock Exchange market.

Table 1: Greek Companies and Acquired Companies

Announcement date	Acquirer Company	Sector	Target Country	Sector	% of acquisition
20/9/2002	Eurodrip	Industrial	Jordan	Industrial	100%
4/6/2002	ETEM	Industrial	Romania	Industrial	20%
2/4/2002	Intracom Constr.	Construction	Bulgaria	Construction	30%
10/05/01	Sidenor	Industrial	Bulgaria	Industrial	75%
25/02/04	Sidenor	Industrial	Bulgaria	Industrial	6%
28/12/2001	TITAN	Industrial	Serbia & Montenegro	Industrial	70%
5/7/2002	TITAN	Industrial	Egypt	Construction	44%
1/12/2003	TITAN	Industrial	Skopje	Industrial	47%
17/07/01	Chalkor	Industrial	Bulgaria	Industrial	93%
19/12/2003	Chalkor	Industrial	Bulgaria	Industrial	7%

Table 1 presents the population of the study, which consists of 10 events of acquisitions of shares that six Greek companies performed.

The aim of this study is to detect and analyse the effects of the Greek firms' acquisition announcement on their stock returns and the asymmetric effect of good and bad news. We use the event study methodology of Brown & Warner (1980, 1985). A great advantage of the OLS estimation technique is that, residuals of a stock sum up to 0 in the estimation period, in order to counteract coefficient α's bias with β's. The estimation period begins 211 prior the announcement day and ends 11 before. The event period starts 10 days prior the announcement and ends 10 days after (-10, +10).

Brown and Warner explain the OLS Market Model (1985) and use it to calculate the Abnormal Returns (AR).

$$AR_{i,t} = R_{it} - (\alpha_i + \beta_i * R_{mt}) \tag{1}$$

where AR_{it} is the abnormal return of firm i on day t, R_{it} is the rate of return for stock i on day t, α_i and β_i are OLS coefficients from the estimation period, and R_{mt} is the market return on day t.

In order to find out the impact the announcements have on the stocks, we calculate the average of the Abnormal Returns (AAR), which implies that particular change. We use the period of -211, +10 to calculate the following:

$$AAR_t = \frac{\sum_{t=-211}^{10} AR_{it}}{n} \tag{2}$$

where AR_{it} is the abnormal return for the i^{th} firm on day t and n is the length of the estimation period. According to the theory, when abnormal performance is spread in a period, that is, not clustered, the best way to calculate AR is CAR. Cumulative Average Abnormal Return (CAAR) is the sum of AAR_t of the firms during the estimation period –211, +10, that is:

$$CAAR_{(-211,+10)} = \sum_{t=-211}^{10} AAR_t \tag{3}$$

The t statistic of the CAAR is used to test the hypothesis whether the AAR on the exact day of the announcement and the CAAR during the estimation period are both zero. Since the event dates spread into periods, we can assume cross sectional independence of the data. The t statistic of AAR is calculated as follows:

The standard deviation of the AR_{it} is found as follows:

$$SD(AR_{it}) = \sqrt{\frac{\sum_{t=-211}^{10} (AR_{it} - A\overline{R}_i)^2}{n-1}} \tag{5}$$

where AR_{it} is the abnormal returns of firm i on day t, $A\overline{R}_i$ is the mean of the abnormal returns of firm i and n is the number of time observations [n=211 + 1 (t=0) + 10=222].
Then, t statistic then is:

$$t = \frac{A\overline{R}_{it}}{S\overline{D}_{A\overline{R}}} \tag{6}$$

where $\overline{SD}_{A\overline{R}}$ is the average standard deviation of the mean abnormal returns in event period calculated as shown in equation (5).

The asymmetric effect of 'good' and 'bad' news on stocks' volatility is an interesting feature, which is captured by the Exponential General Autoregressive Conditional Heteroscedasticity (E-GARCH) model. This particular model allows for negative coefficients, while when using the standard GARCH model, it is necessary to ensure that all of the estimated coefficients are positive. The tendency for volatility to decline when returns rise and to rise when returns fall is called leverage effect. The E-GARCH model allows for the asymmetric effect of good and bad news in the estimation period of an acquisition to take place. We use period -211,-1 of AAR to test this effect.

The form of our E-GARCH model is as follows:

$$AAR_{-211,-1} = \beta_0 + \beta_1 D_1 + \beta_2 D_2 + \varepsilon_t \tag{7}$$

where $AAR_{-211,-1}$ is the estimated average abnormal return of -211,-1 period, D_1 and D_2 are two dummy variables for good and bad news respectively and are:

$$D_1 = \begin{cases} 1 & \text{if } t \in \{-211,-11\} \\ 0 & \text{if } t \in \{-10,-1\} \end{cases} \qquad D_2 = \begin{cases} 0 & \text{if } t \in \{-211,-11\} \\ 1 & \text{if } t \in \{-10,-1\} \end{cases}$$

and

$$\ln(h_t) = \alpha_0 + \alpha_1 \frac{\varepsilon_{t-1}}{h_{t-1}^{0\cdot 5}} + \lambda_1 \left| \frac{\varepsilon_{t-1}}{h_{t-1}^{0\cdot 5}} \right| + \alpha_2 \ln(h_{t-1}) \tag{8}$$

Assumptions of the model are:

$$\varepsilon_t^2 = v_t^2 \,.\, h_t \tag{9}$$

$$E_{t-1} \,.\, \varepsilon_t^2 = h_t \tag{10}$$

$$\text{or } h_t = \alpha_0 + \sum_{i=1}^{q} a_i \varepsilon_{t-i}^2 + \sum_{i=1}^{p} \beta_i h_{t-i} \tag{11}$$

where ε_t^2 is the squared error term, h_t is the conditional variance of ε^2 and $E_{t-1} \,.\, \varepsilon_t^2$ is the lagged expected value of ε_t^2.

Our model permits some coefficients to be negative and the standardised value of ε_{t-1} allows for more natural interpretation of the size and persistence of shocks. If coefficient of $\frac{\varepsilon_{t-1}}{h_{t-1}^{0\cdot 5}}$ is positive (negative), the effect of the shock on the log of the conditional variance is equal to $\alpha_1 + \lambda_1$ $(-\alpha_1 + \lambda_1)$. This is a way of allowing financial leverage effects.

EMPIRICAL RESULTS

Looking at table 2, there is a variation between negative and positive AARs before the announcement, while all the CAARs at the same period are positive (exceptions exist). Negative CAAR seems to cluster between day -140 and -179. On the day of the announcement, t=0, AAR goes slightly above zero, while CAAR are still positive. At day zero, we have 0.4% AARs and 1.3% CAARs.

Both AARs and CAARs on day 0 are statistically insignificant (t value=0.255 and 0.772), which means that we cannot reject the hypothesis that the announcement of the acquisitions does not affect abnormal returns on that day. Similarly, the values of AAR and CAAR on day 0 (0.4% and 1.3% correspondingly) are economically insignificant as well. These findings are in line with other studies of different sectors (i.e. banking and financial sector of industry). Therefore, we can conclude that the announcement of an acquisition does not have a significant impact on the firms' stock values.

The results of the E-GARCH technique are (t-values in parentheses):

$$AAR_{-211,-1} = \underset{(0.2686)}{0.00113.D_1} - \underset{(-0.6459)}{0.0011157.D_2} \tag{12}$$

while the parameters of the conditional heteroscedasticity model are:

$$\ln(h_t) = -10.2471 - 0.028391 \frac{\varepsilon_{t-1}}{h_{t-1}^{0.5}} - 0.64292 \left| \frac{\varepsilon_{t-1}}{h_{t-1}^{0.5}} \right| \tag{13}$$

s.e. (0.11760) (0.10846) (0.17614)

Table 2: AARt and CAARt of event period -211, +10

t	AAR	t value	CAAR	t value
-211	0.00006	0.00338***	0.00006	0.00338***
-200	0.02127	1.24788	0.03140	1.84206
-180	-0.00136	-0.07957*	0.00156	0.09126
-160	-0.00310	-0.18171	-0.03083	-1.80858
-140	0.00656	0.38458	-0.00948	-0.55600
-120	-0.00682	-0.40009	0.00028	0.01671**
-100	0.00678	0.39760	0.04583	2.68814
-80	0.00371	0.21786	0.04049	2.37520
-60	-0.00603	-0.35343	0.00388	0.22771
-40	-0.00145	-0.08479*	0.03230	1.89467
-20	0.00027	0.01570	0.01045	0.61276
-10	0.00247	0.14502	0.02250	1.32001
-5	-0.00222	-0.13015	0.00809	0.47439
-4	0.00263	0.15409	0.01071	0.62848
-3	0.00261	0.15295	0.01332	0.78143
-2	-0.00218	-0.12795	0.01114	0.65349
-1	-0.00233	-0.13674	0.00881	0.51675
0	0.00436	0.25575	0.01317	0.77251
1	-0.00069	-0.04065**	0.01248	0.73186
2	-0.00538	-0.31586	0.00709	0.41600
3	-0.00068	-0.03961**	0.00642	0.37639
4	-0.00309	-0.18145	0.00332	0.19494
5	-0.00258	-0.15138	0.00074	0.04356**
10	0.00394	0.23127	0.00001	0.00059***

This table shows the average and cumulative average abnormal returns around Greek firm mergers.

The log-linear form of the conditional variance's equation allows coefficients to be negative, while a standard GARCH model does not allow for that, as mentioned before. It is clear that bad news have a negative but small effect (- 0.0011157) on the abnormal returns up until day -1, while the presence of good news positively affect AAR in the same period (0.00113). In equation (12), since α_1 is negative, the effect of the shock on the conditional variance h_t is:

$$h_t = - \alpha_1 + \lambda_1 = -0.028391 - 0.64292 = - 0.67131 \qquad (14)$$

that is, shocks on stock prices have a negative effect on the conditional variance.

CONCLUSIONS

This study examines the effect of acquisition announcements on the abnormal returns of six Greek industrial and constructing firms in the time of 2001-2004. We calculated the average and cumulative average abnormal returns using the Market model in combination with the event study methodology. The OLS parameters (α and β) of the model were calculated from the estimation period -211, -11 and were applied for the calculation of AAR and CAAR of period -211, +10. We found the AAR and CAAR for the period -211, +10. Firstly, results show that, the announcement (t = 0) of an acquisition does not significantly affect the AAR and CAAR. Although AAR is positive on day 0, it begins to decline right after that. Apart from that, CAAR is positive between days -120 and way after the announcement day. More generally, it seems that both AAR and CAAR on day t = 0 are statistically and economically insignificant, that is, they do not seem to have a great effect on abnormal returns on that particular day. Secondly, we use the Exponential GARCH model technique to find if there is any correlation between the current return and the future volatility. We include 2 dummy variables in our estimation, namely D_1 and D_2 , which measure the effect of 'good' and 'bad' news on abnormal returns respectively. Results from the E-GARCH model show that the presence of 'good news' positively affects the average abnormal returns in the pre-announcement period -211, -1, while the presence of 'bad news' has a slight negative effect.

We anticipate that the issues addressed in this study will receive further attention by others. We encourage researchers to extend the present study by examining the actual distributions of abnormal return levels across firms or to apply the same methodology by using a different event date, for example, rumours date or outcome date, as t=0.

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