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Editorial Staff: Tele. No: 0092 - 42 - 36560969
Telefax: 0092 - 42 - 36560905
E-mail: nina@lahoreschool.edu.pk, nina_lse@yahoo.com
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Modified Variance Ratio Test for Autocorrelation in the Presence of Heteroskedasticity

Sohail Chand* **and Nuzhat Aftab****

Abstract

Given that autocorrelation tests do not perform well in the presence of heteroskedasticity and in variance-break cases, we present three modified weighted variance ratio tests of autocorrelation. The numerical results show that the proposed tests perform better for small samples. They provide a better approximation of asymptotic distributions and are more powerful when the lag length is mis-specified. The study also applies these tests to data on the daily returns of two companies listed on the Pakistan Stock Exchange.

Keywords: Regression, variance break, wild bootstrap.

JEL classification: C40.

1. Introduction

Testing for autocorrelation in data is an important step in building a linear regression model. While the literature provides many tests for this purpose, the Breusch–Godfrey (BG) Lagrange multiplier (LM) test suggested by Breusch (1978) and Godfrey (1978) is the most popular among practitioners. The objective of these tests of misspecification is to obtain correct inferences about autocorrelation. Ignoring autocorrelation in data can lead to wrong conclusions when testing the significance of the regression parameters of central interest.

Generally, the BG-LM test is used without considering the presence of heteroskedasticity. Several studies investigate the performance of autocorrelation tests in the presence of conditional heteroskedasticity, including Wooldridge (1991), Guo and Phillips (2001) and Mantalos and Shukur (2005), among others. More recently, Jeong and Kang (2012) and Hyun et al. (2010) have used simulations to show that the BG-LM test can generate substantial size distortions in the presence of unconditional heteroskedasticity and misleading results in the case of a variance break.

*Associate Professor, College of Statistical and Actuarial Sciences, University of the Punjab, Pakistan.
** PhD Scholar, College of Statistical and Actuarial Sciences, University of the Punjab, Pakistan.

Shim et al. (2006) propose a modified LM test based on feasible generalized least squares (FGLS) and a revised variance ratio (VR) test for autocorrelation as an alternative to the BG LM test in the presence of variance breaks. Importantly, the FGLS-based LM test is not robust to the type of heteroskedasticity. Moreover, if the heteroskedasticity is not due to a variance break or if no prior information on the breakpoints is available, then the FGLS-based test gives invalid results.

Jeong and Kang (2012) argue that the VR test proposed by Shim et al. (2006), while robust to the type of heteroskedasticity, does not work well for finite samples. They suggest a modified VR test and wild bootstrap version of different VR tests in the presence of unconditional heteroskedasticity. Mun et al. (2014) propose a modified LM test for autocorrelation when there is a break in variance. Their simulation results show that the proposed test has good size and power properties when applied to small samples and is robust to variance-break heteroskedasticity.

This study focuses on the low power of these tests under a mis-specified lag length, especially for small samples. We propose three new VR tests by modifying the weights used in their standard versions. The simulation results show that these modified tests perform reasonably well for both static and dynamic regression models compared to existing tests. Moreover, they are robust to the presence of heteroskedasticity, have good size properties and remain more powerful even in variance-break cases.

The rest of the paper is organized as follows. Section 2 defines the model and VR tests used. The proposed modified VR tests are presented in Section 3. The Monte Carlo simulation results are given in Sections 4 and 5. Section 6 applies these tests to real-life contexts and Section 7 concludes the study.

2. Model and VR Test

We consider the following linear regression model:

$$y_i = X_i^t \beta + u_i, i = 1, 2, \dots, n \tag{1}$$

where y_i is the response variable, X_i is the ith row of the matrix X of order ($n \times (k+1)$), n is the sample size and k is the number of parameters. β is a vector of parameters of order $((k+1) \times 1)$ and $E(u_i) = 0$ for all i. We are interested in whether u_i is serially uncorrelated. The BG-LM test for autocorrelation is reliable when the disturbance term u_i has a constant variance, that is,

homoskedastic errors, but can be misleading in the case of heteroskedastic errors. To consider this scenario, we define the error term u_i in model (1) as

$$u_i = \varepsilon_i \eta_i \tag{2}$$

where ε_i is a heteroskedasticity factor used to create different types of heteroskedasticity, η_i is white noise with a zero mean and unit variance, and $E(\eta_i \eta_j) = 0$ for $i \neq j$. The factor ε_i will result in multiplicative heteroskedasticity if it is a function of a pre-determined variable. In model (2), if we define the factor ε_i in the form of a categorical variable, then it will produce multiple variance-break cases as follows:

$$\varepsilon_i = \sigma_1^2 1[i \leq \tau_1 n] + \sigma_2^2 1[\tau_1 n < i \leq \tau_2 n] + \ldots + \sigma_{s+1}^2 1[i > \tau_s n] \tag{3}$$

where 1[.] is the indicator function and $\tau \in (0, 1)$ is the break timing. Shim et al. (2006) and Hyun et al. (2010) study the simple variance break by considering a single variance breakpoint. In this case, the variance of u_i, that is, σ_i^2, is σ_1^2 for $i \leq \tau_1 n$ and σ_2^2 for $i > \tau_1 n$. Similarly, when we have s breaks, there will be $(s + 1)$ variances in ε_i.

The VR test is designed to identify the presence of serial correlation in a time series by investigating whether the variance of the q-period return is exactly q times larger than the variance of the one-period return. While the literature provides different versions of the VR test, in its simplest form the test compares the variance of the q-period return and q times the variance of the one-period return, that is:

$$VR(q) = \frac{\hat{\sigma}(q)^2}{q\hat{\sigma}(1)^2} - 1 \tag{4}$$

where $q > 1$, $\hat{\sigma}(q)^2 = [\frac{n}{q(n-q+1)(n-q)}] \sum_{i=q}^{n} u_i(q)^2$, $\hat{\sigma}(1)^2 = [\frac{1}{n-k}] \sum_{i=1}^{n} u_i^{\,2}$, $u_i(q) = \sum_{j=0}^{q-1} \hat{u}_{i-j}$ and $\hat{u}_i = y_i - X_i\hat{\beta}$, where $\hat{\beta}$ is the estimator of β. If the errors are serially uncorrelated, then the q-period return variance will be precisely q times the variance of the single-period return. In this case, the value of VR(q) will be near 0.

Lo and MacKinlay (1988) show that the VR(q) given in model (4) can be written as

$$VR(q) = 2\left(1 - \frac{1}{q}\right)\rho(1) + 2\left(1 - \frac{2}{q}\right)\rho(2) + \cdots + 2\rho(q-1) \tag{5}$$

$$VR(q) \approx 2\sum_{h=1}^{q-1}\left(1-\frac{h}{q}\right)\rho(h) \tag{6}$$

where q is the lag length and $\rho(h) = \sum_{i=h+1}^{n}\hat{u}_i\hat{u}_{i-h}/\sum_{i=1}^{n}\hat{u}_i^2$ is the autocorrelation coefficient estimator at lag h. They argue that the VR(q) test has the following limiting distribution under the null hypothesis:

$$\sqrt{nq}\,VR(q) \overset{a}{\sim} N\left(0, \frac{2(2q-1)(q-1)}{3q}\right). \tag{7}$$

Model (5) describes the computation of variance ratios for an aggregate value q. We can see that these are a linear combination of the first $(q-1)$ autocorrelation coefficient estimators with mathematically declining weights. When $q = 2$, the VR(q) value will be approximately the first-order autocorrelation coefficient.

Jeong and Kang (2012) argue that the standard VR(q) test given in models (4) and (6) may be misleading in the presence of negative autocorrelation among the errors. For example, for $q = 3$, the VR(q) approaches 0 if $\Sigma\hat{u}_i\hat{u}_{i-2} = \Sigma\hat{u}_i\hat{u}_{i-1} = \Sigma\hat{u}_{i-2}\hat{u}_{i-1} = 0$, which implies that the errors have no serial correlation. When we have serially correlated errors, the VR(q) statistic will still approach 0, but in this case, $\Sigma\hat{u}_i\hat{u}_{i-2} = -\Sigma\hat{u}_i\hat{u}_{i-1} - \Sigma\hat{u}_{i-2}\hat{u}_{i-1} \neq 0$.

Pagan (1984, 1986) shows that, while the presence of negative autocorrelation might not lead the value of VR(q) towards 0, it can affect the power of the test. To resolve this, Jeong and Kang (2012) suggest using the quadratic version of the VR(q) test, QVR(q), which is defined as:

$$QVR(q) = \sum_{h=1}^{q-1}\left(1-\frac{h}{q}\right)\rho(h)^2 \tag{8}$$

Here, Jeong and Kang (2012) employ the bootstrap distribution, indicating that the asymptotic distribution for it was not derived.

3. Modified VR Tests

This section proposes three new tests: WVR, WQVR and WAVR. The first two are modified versions of the VR and QVR tests. The third test suggests using absolute autocorrelation, given that (as discussed in Section 2) the presence of negative autocorrelation adversely affects the performance of VR tests. Unlike the QVR test, this uses the original magnitude of

autocorrelation while addressing the problem of negative autocorrelation. The new tests are defined as follows:

$$WVR(q) = p\sum_{h=1}^{q-1} W_h \rho(h) \tag{9}$$

$$WQVR(q) = p\sum_{h=1}^{q-1} W_h \rho(h)^2 \tag{10}$$

$$WAVR(q) = p\sum_{h=1}^{q-1} W_h |\rho(h)| \tag{11}$$

where $p = q - 1$, $W_h = \alpha(1-\alpha)^{h-1}$ and $0 < \alpha < 1$. The rationale for using the W_h term is that it enables one to simultaneously control the lag length and assign geometrically decaying weights to the autocorrelations. The parameter α is user-defined but based on empirical evidence, we propose that $\alpha = 0.7$. For these suggested tests, we do not attempt to derive their asymptotic distribution since we employ the wild bootstrap procedure for these statistics.

4. Numerical Results

This section provides numerical results for the empirical size and power of the suggested VR tests, including a comparison with the existing tests. We consider two types of heteroskedasticity: multiplicative heteroskedasticity and the variance-break case. In both cases, we use static *l* and dynamic regression models. In comparing our results with those reported in the literature, we follow the simulation design used by Jeong and Kang (2012).

The static regression model is defined as:

$$y_i = \alpha_0 + \alpha_1 X_i + u_i, i = 1, 2, \dots, n \tag{12}$$

where the predictor $X_i \sim N(0, 1)$ and the values of α_0 and α_1 are set to 1.

The dynamic regression model is given as:

$$y_i = \beta_0 + \beta_1 y_{i-1} + \beta_2 X_i + \beta_3 X_{i-1} + u_i, i = 1, 2, \dots, n \tag{13}$$

The value of β_1 is set to 0.5 to avoid the complexity of nonstationarity, while the remaining parameters are set to 1. To remove the effect of initial values, we have burned out the first 100 simulated observations.

To compute empirical size, we define the serially independent errors of the regression model of multiplicative heteroskedasticity as follows:

$$u_i = \varepsilon_i \eta_i \tag{14}$$

where $\varepsilon_i \sim N(0,1)$, sorted by absolute values, is used to create the heteroskedasticity and $\eta \sim U(-0.5, 0.5)$ is the random error term. Thus, var(u_i) is increasing in i. To compare power, we consider the following seven data-generating processes (DGPs) for the error term:

$$AR(1)+: \quad u_i = 0.7u_{i-1} + \varepsilon_i \eta_i \tag{15}$$

$$AR(1)-: \quad u_i = -0.7u_{i-1} + \varepsilon_i \eta_i \tag{16}$$

$$AR(2): \quad u_i = -0.5u_{i-1} + 0.4u_{i-2} + \varepsilon_i \eta_i \tag{17}$$

$$AR(4): \quad u_i = 0.8u_{i-1} + 0.0u_{i-2} + 0.2u_{i-3} + 0.06u_{i-4} + \varepsilon_i \eta_i \tag{18}$$

$$MA(1)+: \quad u_i = \varepsilon_i \eta_i + 0.7\varepsilon_{i-1}\eta_{i-1} \tag{19}$$

$$MA(1)-: \quad u_i = \varepsilon_i \eta_i - 0.7\varepsilon_{i-1}\eta_{i-1} \tag{20}$$

$$MA(2): \quad u_i = \varepsilon_i \eta_i - 0.5\varepsilon_{i-1}\eta_{i-1} + 0.4\varepsilon_{i-2}\eta_{i-2} \tag{21}$$

The following wild-bootstrap algorithm is used to compute size and power. The only difference between the two is that, for empirical size, the sample is simulated under the null hypothesis, while for power, the simulation is under the alternative hypothesis. In both cases, the model used under the null hypothesis is fitted to the simulated sample. T generally denotes any of the tests, that is, VR, QVR, WVR, WQVR and WAVR.

1. Compute the OLS residuals, $\hat{u}_i = y_i - X_i\hat{\beta}$.
2. Calculate the test (say, T) using the OLS residuals obtained in step 1.
3. Bootstrap y_i (say, y_i^*) using $y_i^* = X_i\hat{\beta} + \{\frac{\hat{u}_i V_i}{\sqrt{(1-h_i)}}\}$, where $h_i = X_i^t(X_t X)^{-1}X_i$ and V_i is a random vector such that $E(V_i) = 0$ and $E(V_1^2) = 1$. Here, V_i is taken from the Rademacher distribution as $\Pr(V_i = 1) = \Pr(V_i = -1) = 0.5$.

4. Repeat steps 1–3 B times (the number of bootstrap samples) to obtain the test for each bootstrap sample, that is, $T_1^*, T_2^*, \ldots, T_B^*$.
5. Compute the $100(1-\alpha)\%$ percentile (say, $T_{1-\alpha}^*$) of the bootstrap distribution obtained in step 4.
6. Repeat steps 1–5 N times. The empirical size and power are computed as $\#(T > T_{1-\alpha}^*)/N$.

Kim (2006) and Jeong and Kang (2012) show that, using the bootstrap method, one can construct an empirical (bootstrap) distribution for the VR test that is robust to the form of heteroskedasticity. They argue that a bootstrap distribution gives more accurate critical values than an asymptotic normal distribution. We consider three choices of sample size – n = 30, 50 and 100 – and four different lag lengths of the autoregressive serial correlation, p = 1, 2, 4, 12. The nominal size of the test is fixed at 5 percent and $q = p + 1$. The number of Monte Carlo runs is set to 1,000, while 500 bootstrap samples are used.

The results for the empirical size of the proposed tests for static and dynamic models are given in Tables 1 and 2, respectively. The static and dynamic models given in (12) and (13) are considered null models. The results given in Table 1 show that, when the null model is static, the size of the proposed test is a very close approximation to the corresponding nominal size. While the standard VR and QVR tests also yield a good size approximation, they deviate from the empirical size the larger the chosen lag. The WVR, WQVR and WAVR tests give better size results and remain stable at varying lag values. Thus, the approximation to the asymptotic distribution of the proposed tests is not affected by the mis-specified lag length.

Table 1: Size comparison for static model with multiplicative heteroskedasticity

	T = 30				T = 50				T = 100			
Test	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**
VR	0.044	0.053	0.062	0.044	0.045	0.043	0.055	0.062	0.055	0.054	0.049	0.074
QVR	0.055	0.050	0.061	0.055	0.043	0.053	0.052	0.050	0.055	0.059	0.056	0.056
WVR	0.053	0.052	0.052	0.048	0.054	0.049	0.057	0.056	0.055	0.044	0.058	0.042
WQVR	0.050	0.048	0.053	0.056	0.049	0.047	0.043	0.046	0.051	0.040	0.056	0.055
WAVR	0.05	0.052	0.054	0.062	0.049	0.039	0.043	0.048	0.051	0.044	0.058	0.051

Note: Nominal size = 5%.
Source: Authors' calculations

Table 2: Size comparison for dynamic model with multiplicative heteroskedasticity

	T = 30				T = 50				T = 100			
Test	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**
VR	0.062	0.055	0.058	0.048	0.048	0.054	0.054	0.053	0.045	0.050	0.055	0.046
QVR	0.043	0.052	0.047	0.048	0.050	0.044	0.054	0.055	0.036	0.049	0.055	0.058
WVR	0.053	0.047	0.054	0.054	0.052	0.044	0.047	0.052	0.050	0.045	0.059	0.055
WQVR	0.040	0.034	0.051	0.051	0.056	0.051	0.038	0.045	0.043	0.054	0.048	0.051
WAVR	0.040	0.036	0.047	0.053	0.056	0.049	0.039	0.047	0.043	0.054	0.048	0.053

Source: Authors' calculations

Table 2 gives the results for the dynamic model defined in (13). In this case, the proposed WVR test shows a closer size approximation than the other tests at different lag values. The numerical results show that the size of the modified VR tests is comparable with that of the classical VR tests as the lag and sample size increases.

The results of empirical power against various alternatives are given in Tables 3 and 4. The sample is generated under seven different models: AR(1)+, AR(1)−, AR(2), AR(4), MA(1)+, MA(1)− and MA(2), as defined in equations (15) to (21). The results show that the proposed tests generally have a higher empirical power than the existing tests, outperforming the latter for all considered choices of p. For all considered alternative DGPs, we find that the power of the VR and QVR tests decreases as the lag increases.

Table 3: Empirical power for static model of multiplicative heteroskedasticity

		T = 30				T = 50				T = 100			
	Test	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**
	VR	0.926	0.927	0.826	0.608	0.992	0.988	0.959	0.785	1	1	1	0.951
	QVR	0.835	0.85	0.805	0.716	0.986	0.976	0.963	0.939	0.999	1	1	1
AR(1)+	WVR	0.825	0.894	0.906	0.906	0.99	0.986	0.981	0.982	1	1	1	0.999
	WQVR	0.825	0.827	0.844	0.838	0.975	0.98	0.974	0.974	1	1	0.999	0.999
	WAVR	0.825	0.827	0.822	0.822	0.975	0.976	0.97	0.969	1	1	0.999	0.999
	VR	0.665	0.368	0.234	0.168	0.851	0.591	0.378	0.164	0.99	0.891	0.71	0.289
	QVR	0.665	0.663	0.607	0.511	0.851	0.844	0.801	0.722	0.99	0.983	0.974	0.923
AR(1)-	WVR	0.916	0.935	0.921	0.913	0.991	0.997	0.991	0.989	1	1	1	1
	WQVR	0.916	0.925	0.917	0.904	0.991	0.994	0.989	0.985	1	1	1	1
	WAVR	0.916	0.907	0.91	0.9	0.991	0.994	0.987	0.985	1	1	1	1
	VR	0.902	0.297	0.219	0.233	0.986	0.6	0.386	0.217	0.999	0.911	0.709	0.296
	QVR	0.902	0.946	0.946	0.93	0.986	0.988	0.991	0.99	0.999	1	1	1
AR(2)	WVR	0.937	0.955	0.954	0.955	0.982	0.987	0.992	0.996	1	1	1	1
	WQVR	0.937	0.898	0.891	0.889	0.982	0.963	0.971	0.961	1	0.999	0.997	0.996
	WAVR	0.937	0.862	0.879	0.865	0.982	0.954	0.967	0.956	1	1	0.997	0.994
	VR	0.925	0.919	0.786	0.456	0.995	0.996	0.958	0.677	1	1	0.988	0.884
	QVR	0.925	0.915	0.890	0.831	0.995	0.996	0.992	0.974	1	1	1	1
AR(4)	WVR	0.960	0.950	0.943	0.949	0.999	0.995	0.993	0.996	1	1	1	1
	WQVR	0.987	0.907	0.905	0.921	0.997	0.995	0.993	0.993	1	1	1	1
	WAVR	0.987	0.900	0.902	0.908	0.997	0.993	0.990	0.992	1	1	1	1
	VR	0.794	0.562	0.404	0.234	0.928	0.756	0.527	0.283	0.998	0.923	0.715	0.414
	QVR	0.575	0.597	0.513	0.352	0.835	0.874	0.823	0.677	0.987	0.994	0.988	0.943
MA(1)+	WVR	0.756	0.858	0.871	0.876	0.939	0.970	0.985	0.979	0.999	1	1	1
	WQVR	0.535	0.543	0.561	0.551	0.84	0.797	0.826	0.826	0.999	0.973	0.975	0.968
	WAVR	0.535	0.539	0.555	0.55	0.84	0.769	0.795	0.807	0.999	0.967	0.965	0.955
	VR	0.458	0.411	0.312	0.215	0.722	0.707	0.584	0.279	0.934	0.941	0.9	0.484
	QVR	0.458	0.451	0.378	0.253	0.722	0.69	0.639	0.472	0.934	0.927	0.898	0.767
MA(1)-	WVR	0.658	0.552	0.526	0.531	0.909	0.779	0.751	0.759	0.996	0.94	0.938	0.945
	WQVR	0.658	0.662	0.662	0.652	0.909	0.873	0.848	0.884	0.996	0.981	0.987	0.99
	WAVR	0.658	0.653	0.664	0.645	0.909	0.852	0.835	0.867	0.996	0.97	0.98	0.987
	VR	0.705	0.356	0.157	0.104	0.871	0.530	0.250	0.098	0.977	0.823	0.441	0.113
	QVR	0.705	0.699	0.657	0.547	0.871	0.870	0.875	0.788	0.977	0.992	0.988	0.969
MA(2)	WVR	0.668	0.725	0.699	0.693	0.865	0.881	0.884	0.868	0.972	0.987	0.988	0.979
	WQVR	0.668	0.695	0.664	0.670	0.865	0.849	0.860	0.859	0.972	0.962	0.976	0.974
	WAVR	0.668	0.669	0.648	0.640	0.865	0.830	0.841	0.838	0.972	0.949	0.969	0.969

Source: Authors′ calculations

Table 4: Empirical power for dynamic model of multiplicative heteroskedasticity

		T = 30				T = 50				T = 100			
	Test	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**
	VR	0.841	0.830	0.700	0.514	0.981	0.979	0.919	0.698	0.999	1	0.997	0.919
	QVR	0.647	0.621	0.563	0.507	0.926	0.924	0.899	0.821	0.999	1	0.999	0.989
AR(1)+	WVR	0.849	0.806	0.803	0.815	0.982	0.959	0.965	0.973	1	0.999	1	1
	WQVR	0.618	0.640	0.641	0.636	0.933	0.932	0.931	0.934	1	0.997	1	1
	WAVR	0.618	0.609	0.628	0.618	0.933	0.918	0.917	0.913	1	0.996	0.999	1
	VR	0.848	0.292	0.168	0.132	0.990	0.680	0.434	0.184	1	0.996	0.825	0.378
	QVR	0.848	0.870	0.815	0.734	0.990	0.980	0.985	0.942	1	1	1	0.998
AR(1)-	WVR	0.865	0.872	0.866	0.869	0.984	0.983	0.985	0.981	1	1	1	1
	WQVR	0.865	0.840	0.823	0.833	0.984	0.980	0.981	0.979	1	1	1	1
	WAVR	0.865	0.801	0.795	0.803	0.984	0.980	0.979	0.972	1	1	1	1
	VR	0.833	0.058	0.053	0.067	0.973	0.264	0.097	0.06	0.999	0.738	0.379	0.119
	QVR	0.833	0.909	0.912	0.898	0.973	0.986	0.989	0.987	0.999	1	1	1
AR(2)	WVR	0.861	0.928	0.92	0.933	0.971	0.99	0.991	0.993	1	1	1	1
	WQVR	0.861	0.759	0.777	0.763	0.971	0.947	0.947	0.947	1	0.996	0.998	0.997
	WAVR	0.861	0.69	0.739	0.727	0.971	0.932	0.94	0.937	1	0.996	0.999	0.998
	VR	0.924	0.912	0.732	0.438	0.996	0.994	0.939	0.614	1	1	0.997	
	QVR	0.776	0.742	0.678	0.626	0.983	0.997	0.971	0.923	1	1	1	
AR(4)	WVR	0.927	0.906	0.874	0.893	0.993	0.988	0.984	0.991	1	0.999	1	1
	WQVR	0.764	0.798	0.748	0.791	0.977	0.979	0.974	0.982	1	0.999	1	1
	WAVR	0.764	0.771	0.737	0.768	0.977	0.973	0.968	0.976	1	0.999	1	1
	VR	0.771	0.489	0.337	0.219	0.9	0.687	0.467	0.265	0.996	0.989	0.641	0.373
	QVR	0.369	0.392	0.342	0.281	0.746	0.799	0.744	0.58	0.98	0.992	0.979	0.916
MA(1)+	WVR	0.969	0.814	0.812	0.813	0.926	0.975	0.968	0.962	0.996	0.997	1	1
	WQVR	0.383	0.372	0.398	0.397	0.752	0.74	0.728	0.744	0.977	0.955	0.956	0.965
	WAVR	0.383	0.364	0.397	0.39	0.752	0.717	0.716	0.715	0.977	0.943	0.943	0.949
	VR	0.501	0.484	0.336	0.221	0.816	0.848	0.727	0.397	0.988	0.992	0.987	0.788
	QVR	0.501	0.444	0.403	0.297	0.816	0.831	0.75	0.569	0.988	0.994	0.984	0.914
MA(1)-	WVR	0.470	0.441	0.434	0.400	0.828	0.678	0.689	0.679	0.983	0.922	0.930	0.928
	WQVR	0.470	0.503	0.521	0.505	0.828	0.798	0.818	0.794	0.983	0.961	0.979	0.984
	WAVR	0.470	0.504	0.521	0.515	0.828	0.789	0.809	0.785	0.983	0.951	0.974	0.974
	VR	0.589	0.235	0.085	0.084	0.821	0.438	0.167	0.069	0.985	0.730	0.34	0.078
	QVR	0.589	0.571	0.567	0.438	0.821	0.826	0.833	0.703	0.985	0.973	0.979	0.952
MA(2)	WVR	0.604	0.657	0.607	0.585	0.830	0.872	0.846	0.873	0.986	0.985	0.981	0.975
	WQVR	0.604	0.590	0.548	0.558	0.830	0.822	0.810	0.826	0.986	0.959	0.97	0.951
	WAVR	0.604	0.541	0.521	0.536	0.830	0.801	0.776	0.797	0.986	0.946	0.953	0.941

Source: Authors' calculations

Table 4 gives the empirical power results for the dynamic model of multiplicative heteroskedasticity. These are generally similar to the results for the static model. The power of the WVR test is higher, especially for small samples. In the dynamic model, as the lag length increases, the power

of the VR and QVR tests decreases even when the lag length is properly specified. This decrease is more evident in the case of negative autocorrelation, for example, for MA(1)− and AR(1)−. The VR test becomes almost powerless for larger lags in the MA(2) and AR(2) models. Again, the proposed tests indicate good power except in some cases, for instance, MA(1)− for small samples.

5. Variance-Break Case

This section considers variance-break heteroskedasticity. Following Jeong and Kang (2012), the heteroskedastic factor ε_i is defined as a categorical variable as follows:

$$\varepsilon_i^2 = \sigma_1^2 1[i \le \tau_1 n] + \sigma_2^2 1[\tau_1 n < i \le \tau_2 n] + \sigma_3^2 1[i > \tau_2 n]. \tag{24}$$

We have specified $\sigma_1^2 = 1$, $\sigma_2^2 = 4$, $\sigma_3^2 = 9$, $\tau_1 = 0.3$ and $\tau_2 = 0.7$ in this study. Except for the change in the type of heteroskedasticity, the static and dynamic DGPs used are similar to those employed in the previous section.

Table 5 gives the empirical size of the considered VR tests for the static model, using a two-point variance break. The results show that the proposed tests are comparable in terms of empirical size and that the size of the VR test is adversely affected by the presence of a break in variance, especially when the sample is small, that is, for n = 30 and n = 50. The results for the dynamic model with a two-point variance break, given in Table 6, are not much different from the results for the static model. The proposed tests yield a good size approximation for both models in the variance-break case.

Table 5: Empirical size under static model with variance break

	T = 30				T = 50				T = 100			
Test	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**
VR	0.048	0.063	0.059	0.059	0.054	0.059	0.056	0.06	0.056	0.041	0.041	0.048
QVR	0.057	0.056	0.049	0.052	0.06	0.049	0.060	0.048	0.047	0.051	0.044	0.053
WVR	0.055	0.054	0.06	0.054	0.058	0.058	0.060	0.048	0.056	0.046	0.047	0.057
WQVR	0.048	0.051	0.055	0.049	0.058	0.047	0.055	0.057	0.050	0.053	0.052	0.052
WAVR	0.048	0.051	0.052	0.055	0.058	0.044	0.052	0.059	0.050	0.053	0.049	0.052

Source: Authors′ calculations

Table 6: Empirical size under dynamic model with variance break

	T = 30				T = 50				T = 100			
Test	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**
VR	0.060	0.059	0.057	0.06	0.065	0.049	0.056	0.054	0.049	0.045	0.051	0.051
QVR	0.042	0.046	0.048	0.051	0.05	0.048	0.054	0.047	0.038	0.044	0.045	0.055
WVR	0.042	0.044	0.057	0.034	0.05	0.048	0.052	0.047	0.04	0.048	0.04	0.046
WQVR	0.042	0.035	0.04	0.045	0.05	0.043	0.049	0.038	0.038	0.041	0.033	0.04
WAVR	0.042	0.033	0.039	0.043	0.05	0.04	0.045	0.041	0.038	0.041	0.034	0.038

Source: Authors' calculations

The empirical power results of the variance-break case for the five tests are given in Tables 7 and 8 and lead to the same conclusion as for the multiplicative heteroskedasticity case. The tests' good performance is, therefore, unaffected by the change in type of heteroskedasticity.

Table 7: Empirical power under static model with variance-break heteroskedasticity

		T = 30				T = 50				T = 100			
	Test	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**
AR(1)+	VR	0.992	0.896	0.818	0.582	0.994	0.987	0.955	0.803	1	1	1	1
	QVR	0.825	0.794	0.744	0.685	0.973	0.971	0.956	0.936	1	1	1	1
	WVR	0.922	0.626	0.601	0.63	0.994	0.982	0.985	0.985	1	1	1	1
	WQVR	0.825	0.799	0.779	0.786	0.973	0.975	0.97	0.973	1	1	1	1
	WAVR	0.825	0.777	0.761	0.779	0.973	0.962	0.965	0.968	1	1	1	1
AR(1)-	VR	0.859	0.406	0.271	0.25	0.99	0.764	0.542	0.315	1	0.995	0.943	0.611
	QVR	0.859	0.882	0.844	0.771	0.99	0.987	0.98	0.96	1	1	1	1
	WVR	0.859	0.894	0.898	0.876	0.99	0.987	0.984	0.986	1	1	1	1
	WQVR	0.859	0.876	0.889	0.879	0.99	0.987	0.987	0.985	1	1	1	1
	WAVR	0.859	0.857	0.875	0.864	0.99	0.987	0.984	0.98	1	1	1	1
AR(4)	VR	0.955	0.93	0.834	0.515	0.998	0.998	0.949	0.671	1	1	1	0.87
	QVR	0.877	0.876	0.852	0.771	0.993	0.992	0.98	0.982	1	1	1	1
	WVR	0.877	0.736	0.721	0.722	0.993	0.962	0.961	0.96	1	1	1	1
	WQVR	0.877	0.878	0.879	0.877	0.993	0.987	0.985	0.987	1	1	1	1
	WAVR	0.877	0.873	0.856	0.859	0.993	0.986	0.983	0.986	1	1	1	1
MA(1)+	VR	0.701	0.507	0.359	0.235	0.923	0.701	0.493	0.294	1	0.953	0.727	0.441
	QVR	0.473	0.491	0.427	0.356	0.83	0.855	0.776	0.597	0.997	0.998	0.996	0.959
	WVR	0.701	0.835	0.835	0.856	0.923	0.973	0.981	0.97	1	1	1	1
	WQVR	0.476	0.504	0.503	0.528	0.83	0.794	0.792	0.798	0.997	0.982	0.985	0.984
	WAVR	0.473	0.502	0.503	0.524	0.83	0.768	0.778	0.775	0.997	0.979	0.978	0.976
MA(1)-	VR	0.581	0.56	0.461	0.365	0.827	0.607	0.444	0.24	0.994	0.885	0.678	0.962
	QVR	0.581	0.552	0.490	0.36	0.827	0.874	0.791	0.602	0.994	0.998	0.997	0.961
	WVR	0.581	0.463	0.481	0.474	0.827	0.891	0.885	0.879	0.994	0.998	0.998	0.949
	WQVR	0.581	0.593	0.609	0.59	0.827	0.789	0.812	0.787	0.994	0.989	0.985	0.984
	WAVR	0.581	0.578	0.614	0.583	0.827	0.762	0.789	0.772	0.994	0.985	0.978	0.979
MA(2)	VR	0.61	0.289	0.109	0.111	0.842	0.504	0.201	0.085	0.98	0.823	0.462	0.123
	QVR	0.61	0.642	0.594	0.491	0.842	0.851	0.824	0.726	0.98	0.991	0.988	0.967
	WVR	0.61	0.673	0.651	0.624	0.842	0.87	0.854	0.857	0.98	0.992	0.987	0.986
	WQVR	0.61	0.63	0.632	0.612	0.842	0.819	0.838	0.793	0.98	0.975	0.987	0.974
	WAVR	0.61	0.597	0.6	0.593	0.842	0.803	0.798	0.767	0.98	0.968	0.974	0.966
AR(2)	VR	0.85	0.145	0.11	0.143	0.977	0.482	0.249	0.175	1	0.908	0.692	0.277
	QVR	0.85	0.887	0.887	0.858	0.977	0.99	0.981	0.975	1	1	1	1
	WVR	0.85	0.902	0.907	0.912	0.977	0.994	0.986	0.986	1	1	1	1
	WQVR	0.85	0.812	0.827	0.832	0.977	0.962	0.956	0.958	1	1	1	0.998
	WAVR	0.85	0.777	0.801	0.796	0.977	0.955	0.946	0.952	1	1	1	0.998

Source: Authors′ calculations

Table 8: Empirical power under dynamic model with variance-break heteroskedasticity

		T = 30				T = 50				T = 100			
	Test	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**	**p = 1**	**p = 2**	**p = 4**	**p = 12**
AR(1)+	VR	0.825	0.779	0.685	0.456	0.974	0.954	0.906	0.633	1	0.999	0.996	0.889
	QVR	0.572	0.546	0.501	0.409	0.897	0.88	0.854	0.742	0.998	0.997	0.998	0.983
	WVR	0.828	0.769	0.772	0.773	0.974	0.931	0.941	0.944	0.998	0.997	0.998	0.999
	WQVR	0.572	0.55	0.548	0.562	0.897	0.878	0.887	0.861	0.998	0.997	0.998	0.999
	WAVR	0.572	0.539	0.529	0.536	0.897	0.866	0.872	0.854	0.998	0.996	0.998	0.999
AR(1)-	VR	0.777	0.172	0.114	0.098	0.97	0.453	0.231	0.095	1	0.881	0.602	0.164
	QVR	0.777	0.759	0.737	0.629	0.97	0.97	0.961	0.924	1	1	1	1
	WVR	0.777	0.793	0.807	0.811	0.97	0.974	0.973	0.981	1	1	1	1
	WQVR	0.777	0.707	0.728	0.723	0.97	0.953	0.962	0.966	1	1	1	1
	WAVR	0.777	0.662	0.702	0.673	0.97	0.937	0.953	0.948	1	1	1	1
AR(4)	VR	0.894	0.885	0.711	0.400	0.991	0.989	0.907	0.525	1	1	0.992	0.767
	QVR	0.698	0.691	0.635	0.548	0.964	0.963	0.932	0.853	1	1	1	1
	WVR	0.698	0.851	0.846	0.854	0.964	0.978	0.972	0.848	1	1	1	1
	WQVR	0.698	0.704	0.701	0.692	0.964	0.959	0.951	0.931	1	1	1	1
	WAVR	0.698	0.687	0.691	0.676	0.964	0.952	0.941	0.925	1	0.999	0.999	1
MA(1)+	VR	0.651	0.451	0.306	0.203	0.907	0.591	0.383	0.209	0.996	0.906	0.614	0.296
	QVR	0.348	0.345	0.314	0.263	0.700	0.724	0.691	0.483	0.983	0.993	0.986	0.93
	WVR	0.651	0.774	0.805	0.779	0.907	0.947	0.968	0.972	0.996	1	1	1
	WQVR	0.348	0.349	0.365	0.364	0.700	0.642	0.65	0.641	0.983	0.957	0.96	0.963
	WAVR	0.348	0.351	0.365	0.357	0.700	0.619	0.628	0.621	0.983	0.952	0.943	0.945
MA(1)-	VR	0.42	0.353	0.237	0.164	0.742	0.73	0.599	0.302	0.987	0.995	0.982	0.7
	QVR	0.42	0.386	0.322	0.253	0.742	0.716	0.633	0.426	0.987	0.992	0.976	0.888
	WVR	0.42	0.358	0.361	0.360	0.742	0.593	0.592	0.574	0.987	0.927	0.91	0.91
	WQVR	0.42	0.405	0.439	0.415	0.742	0.72	0.723	0.709	0.987	0.972	0.972	0.966
	WAVR	0.42	0.412	0.433	0.415	0.742	0.706	0.702	0.691	0.987	0.964	0.957	0.957
MA(2)	VR	0.51	0.177	0.059	0.062	0.773	0.338	0.092	0.081	0.977	0.653	0.205	0.047
	QVR	0.51	0.526	0.505	0.382	0.773	0.786	0.778	0.743	0.977	0.984	0.989	0.954
	WVR	0.51	0.551	0.561	0.551	0.773	0.812	0.801	0.784	0.977	0.984	0.985	0.975
	WQVR	0.51	0.483	0.475	0.478	0.773	0.743	0.759	0.751	0.977	0.957	0.964	0.954
	WAVR	0.51	0.446	0.445	0.443	0.773	0.709	0.717	0.696	0.977	0.942	0.951	0.938
AR(2)	VR	0.766	0.04	0.026	0.057	0.947	0.073	0.033	0.039	1	0.259	0.059	0.02
	QVR	0.766	0.85	0.831	0.826	0.947	0.984	0.988	0.971	1	1	1	1
	WVR	0.766	0.856	0.861	0.882	0.947	0.986	0.992	0.983	1	1	1	1
	WQVR	0.766	0.61	0.561	0.64	0.947	0.829	0.844	0.845	1	0.965	0.97	0.974
	WAVR	0.766	0.511	0.508	0.573	0.947	0.792	0.837	0.839	1	0.974	0.98	0.986

Source: Authors′ calculations

6. Real-Life Applications

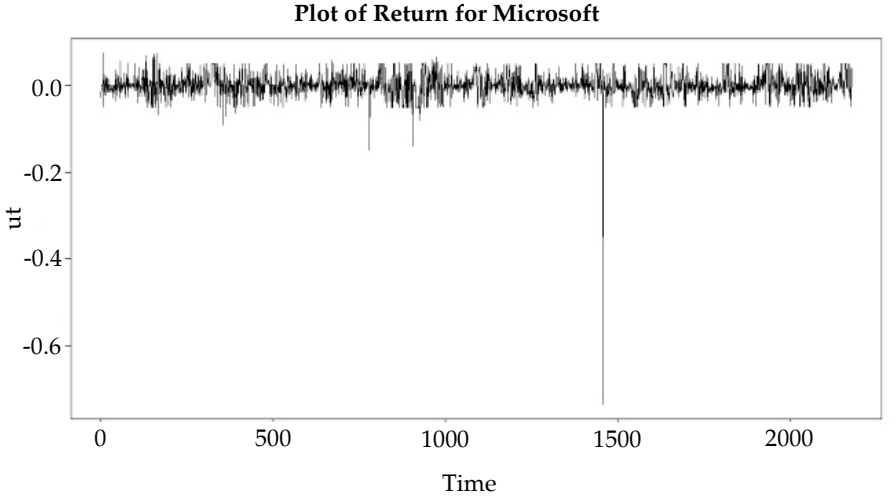
This section applies the classical and proposed VR tests to daily returns data for two companies listed on the Pakistan Stock Exchange and

to the Microsoft Corporation. As Yasmin (2009) points out, heteroskedasticity is very common in financial data series. The data used here represents the daily stock returns of Dawood Hercules (DAWH) and Kohinoor Textile Mills Limited (KTML) for the period 1 July 2004 to 28 February 2014 as well as the daily stock returns of Microsoft for the period May 2000 to May 2017.

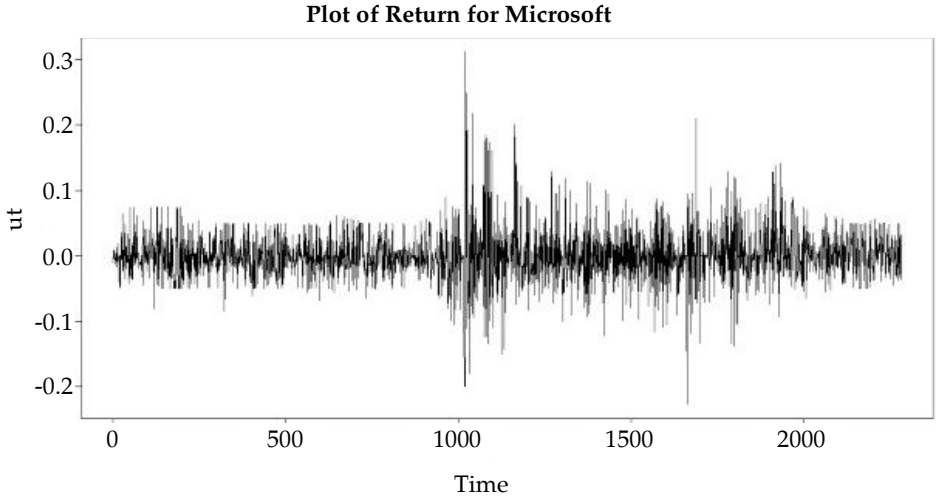
We test the series for the presence of conditional heteroskedasticity, using Engle's (1982) LM test. As Figure 1 shows, the daily returns series for DAWH does not exhibit heteroskedasticity (LM = 0.0717, p-value = 1). The KTML and Microsoft series, however, both indicate the presence of heteroskedasticity (LM = 479.9, p-value = 0.000; and LM = 261.27, p-value = 0.000, respectively).

Figure 1: Daily returns data for DAWH, KTML and Microsoft

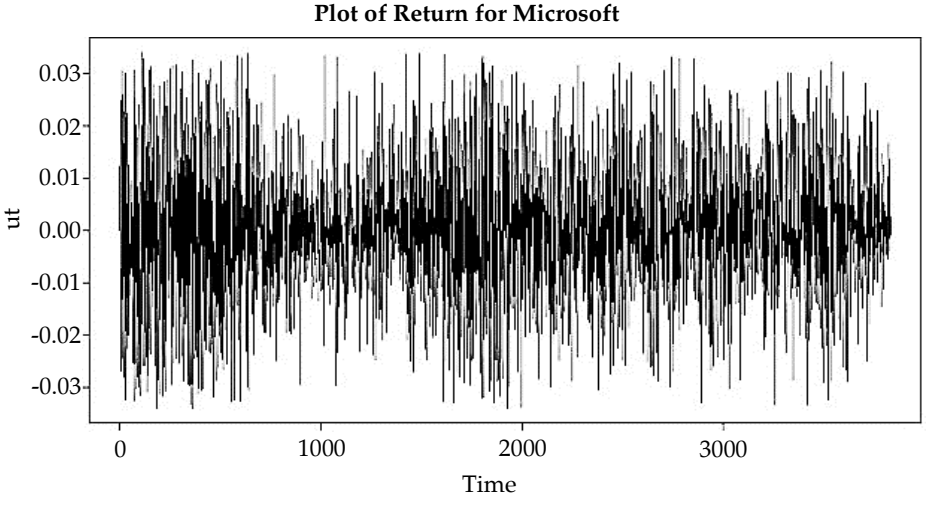
(a) Homoskedasticity (DAWH)



(b) Variance break (KTML)



(c) Multiplicative heteroskedasticity (Microsoft)



The results of the VR tests are given in Table 9. The DAWH return series does not exhibit any conditional heteroskedasticity. All the tests are significant, indicating the presence of autocorrelation in the data even at a 1 percent level of significance for all considered choices of p. In the variance-break case – for the KTML returns – the VR and QVR tests are not significant for any lag value, even at a 10 percent level of significance. The modified QVR and AVR tests show very strong levels of rejection.

When the heteroskedasticity is multiplicative, the VR and WVR tests yield insignificant results at all lag values, while the QVR, WQVR and WAVR tests are significant at lags 1, 2 and 4 at the 5 percent level. The behavior of the proposed tests is consistent with the change in lag value, although the critical region for QVR changes with the lag value. The proposed tests yield significant results for all values of p at the 10 and 5 percent levels of significance and present fairly strong evidence against the null hypothesis.

Table 9: VR tests for daily returns data

p	VR	p value	QVR	p value	WVR	p value	WAVR	p value	WQVR	p value
DAWH										
1	0.096***	0.0000	0.0046***	0.0000	0.0672***	0.0000	0.0672***	0.0000	0.0064***	0.0000
2	0.1653***	0.0000	0.0072***	0.0000	0.1108***	0.0000	0.0018***	0.0000	0.0116***	0.0000
4	0.2447***	0.0000	0.0095***	0.0010	0.2267***	0.0000	0.2267***	0.0000	0.0233***	0.0000
12	0.4199***	0.0000	0.0132***	0.0050	0.6808***	0.0020	0.6808***	0.0010	0.0699***	0.0010
KTML										
1	-0.0345	0.9502	0.0006*	0.0938	-0.0242	0.9502	0.0242*	0.0938	0.0006*	0.0938
2	-0.0451	0.9222	0.0008	0.1634	-0.0489	0.945	0.0477*	0.0818	0.0017*	0.0646
4	-0.0294	0.7268	0.0013	0.2046	-0.0906	0.9296	0.1026*	0.0768	0.0035**	0.0561
12	-0.0994	0.871	0.0038	0.1548	-0.2725	0.9286	0.3085*	0.0802	0.0107**	0.0578
Microsoft										
1	-0.0365	0.9884	0.0007**	0.0254	-0.0255	0.9884	0.0255**	0.0254	0.0009**	0.0254
2	-0.053	0.9846	0.0009**	0.0422	-0.0540	0.9870	0.0540**	0.0408	0.0019**	0.0320
4	-0.078	0.9896	0.0012*	0.0676	-0.1117	0.9926	0.1119**	0.0342	0.0038**	0.0288
12	-0.1475	0.9920	0.0034**	0.021	-0.3376	0.9902	0.3385**	0.0338	0.0115**	0.0262

Note: * = significant at 10% level, ** = significant at 5% level, *** = significant at 1% level.

7. Conclusion

The performance of autocorrelation tests is adversely affected if the errors do not have a constant variance. In this context, we propose three alternative VR tests that yield very good size properties for both small and large samples and are not heavily dependent on the specified lag. These tests retain their power across static and dynamic models regardless of whether the heteroskedasticity is multiplicative or if there is a variance break.

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The Paradigm Shift in the Pakistan Stock Exchange's Financial Integration Post-FTA and CPEC

Abdul Wahid* **and Muhammad Zubair Mumtaz****

Abstract

This paper examines whether regional connectivity causes return and volatility spillovers and the co-movement of stock exchanges to shift from international to regional markets. Using the China-Pakistan free trade agreement (FTA) of 2006 and the China-Pakistan Economic Corridor (CPEC) agreement to represent events of regional connectivity, we test this proposition based on data for two regional stock exchanges (the Pakistan Stock Exchange and Shenzhen Stock Exchange) and two global markets (the FTSE 100 and Nasdaq). We divide the convergence and co-integration of the stock markets into three phases: overall sample (2001–17), pre-FTA and post-FTA, and pre-CPEC and post-CPEC. Applying a GARCH (1, 1) model, co-integration, Granger causality and seasonality, we find that regional connectivity causes return and volatility spillovers and co-movements in the Pakistan Stock Exchange to shift from international markets to regional markets.

Keywords: Financial integration, China-Pakistan economic corridor, interdependency, co-integration, free trade agreement, regional connectivity, Pakistan.

JEL classification: C22, E44, F15, F21, R58.

1. Introduction

Globalization generally develops around the paradigm shift from localization to internationalization, which promotes the co-integration and co-movement of stock exchanges with global markets (Zhang et al., 2015). This phenomenon is based on three pillars – integration, convergence and cohesiveness (Barro, 2016) – which enable greater economic liberalization (Alikhanov, 2013), interrelated synchronization and the financial integration of full-fledged local stock exchanges with international markets (Patel, 2016).

* Lecturer, Department of Governance and Public Policy, Faculty of Management Sciences, National University of Modern Languages, Islamabad, Pakistan.
** Assistant Professor, School of Social Sciences and Humanities, National University of Science and Technology, Islamabad, Pakistan.

This has garnered considerable research interest with respect to co-movements and spillover effects in stock markets. Earlier studies have examined, for example, how the change in volatility and returns of one market can influence the volatility and returns of other markets, that is, how one follows the movements of the other.

The literature looks at aspects of financial integration such as the interdependency of stock exchanges (Rajhans & Jain, 2015), the trickledown effects of integration (Morana & Beltratti, 2008), the co-movement of prices, returns and volatilities (Sehgal et al., 2015), and correlations between developed and developing country markets (Wang & Liu, 2016). Studies on financial integration provide investors with useful insights into international market trends and their trickledown influence over regional markets (Carsamer, 2016). This creates opportunities for international investors to develop stock portfolios for emerging economies (Bhar & Nikolova, 2007; Sehgal et al., 2015).

After the financial crisis of 2007/08, regional connectivity, partnerships and integration have drawn greater interest. Critics of financial integration argue that it leads to the creation of market bubbles (Hatipoglu & Uyar, 2012) and that the high interdependency of markets increases the risk of financial crisis (Morana & Beltratti, 2008), such that highly volatile, risky markets adversely affect stable, emerging ones (Li & Majerowska, 2008). Other studies have focused on stock exchange behavior in integrated systems (Gunasinghe, 2005) and the co-movement of multiple stock exchanges in the long and short runs (Li & Majerowska, 2008; Masih & Masih, 2001; Patel & Patel, 2012) post-globalization and financial liberalism. Joshi (2011), for instance, focuses on how stock exchanges are integrated, while Mulyadi and Anwar (2012) assess to what extent they affect each other in terms of mean or volatility spillovers and over what period.

This debate is rooted in different aspects of globalization, including international partnerships, financial integration and homogeneity of operations (Sial et al., 2015). While globalization is responsible for co-movement and financial integration, countries' economic relationships and their control over other stock exchanges (foreign shareholding) can also be considered major contributors to movements in financial markets (Taneja, 2012). The linkages that develop through these regional interests, strategic partnerships and stock market movements thus represent an important, if under-researched, part of the literature, to which this study contributes.

We focus on the Pakistan Stock Exchange (PSX), which was initially more dependent on international stock markets. In 2006, Pakistan and China signed a mutual free trade agreement (FTA) to promote bilateral trade. Economic cooperation between the two countries grew over time and the China-Pakistan Economic Corridor (CPEC) agreement was established in 2015, with an estimated cost of US$57 billion. Post-CPEC, a Chinese consortium acquired 40 percent of its shares after the PSX was demutualized (Hussain, 2016) and the latter appeared to converge from the global market to the regional market.[1]

Accordingly, we draw on convergence theory, which states that economic and strategic partnerships create a similar culture and business trends. In this case, the entry of Chinese investors through portfolio and foreign direct investment as well as companies' cross-listing has contributed to financial integration, economic convergence, capital market growth and economic development in the region (Hussain, 2016). These economic transactions have enabled Chinese investors to participate in the PSX's decision-making process, in turn, causing it to converge from Western markets to regional markets.

We add to the literature on the co-movement of stock exchanges and return and volatility spillovers by showing how economic and strategic partnerships can trigger movements in a country's capital market toward its partner. Specifically, we ask (i) how the returns and volatility of one country's stock market can influence its economic partner's returns and volatility, and (ii) how such partnerships can lead to co-movements in their stock exchanges.

2. Empirical Evidence of Co-movements and Spillovers

In a recent study, Jebran and Iqbal (2016) examine volatility spillovers between stock markets and foreign exchange markets in selected Asian economies, including Pakistan, India, Sri Lanka, China, Hong Kong and Japan. They present evidence of a bidirectional, asymmetric volatility spillover effect between the stock markets and foreign exchange markets of Pakistan, China, Hong Kong and Sri Lanka. The authors also explore the unidirectional transmission of volatility from the Indian stock market to the

[1] The PSX sold 40 percent of its shares to a Chinese consortium comprising the China Financial Futures Exchange Company Limited, the Shanghai Stock Exchange and Shenzhen Stock Exchange. Valued at PRs8.96 billion (roughly US$85 million), this transaction is expected to attract foreign investment, skills, technological assistance and new product development.

foreign exchange market. In Japan's case, however, there is little evidence of volatility transmission between the two markets.

Peng et al. (2017) investigate the return spillover between Taiwan's TAIEX and the Nikkei, and point to persistent (long-term) co-integration between the two indices. Examining the cross-market effect, they demonstrate how past returns on the Nikkei influence present returns on the TAIEX, although not the other way around. The study reveals the presence of an inner market leverage effect, a negative cross-market volatility spillover effect and a common price-leading effect. Looking at the two indices' relative asymmetry, the authors find that both stock markets are more sensitive to falling than rising patterns in each other.

Liow (2015) examines the contingent volatility spillover effect among five asset classes (public real estate, stocks, bonds, money and currency) locally and internationally in the G7 countries for the period 1997–2013. The author finds that persistent volatility is predominant in all domestic asset markets. The results also indicate an association between the spillover cycle co-movements of domestic business cycle fluctuations and the asset market return volatility cycle. This provides fresh insight into cross-asset and cross-market volatility for national bank strategy makers involved in risk management and portfolio diversification for international investors.

Tanizaki and Hamori (2009) use price level and volatility to analyze spillover effects among Japan, the UK and US. They gauge the impact of return and volatility spillovers among the stock market, credit default swap (CDS) market and foreign exchange market in Korea, the US and Japan. The authors present evidence of a critical return and volatility spillover effect between the Korean CDS and stock markets. Taly (2015) points to a return spillover effect in foreign exchange markets and from the US stock market to the Korean stock market, as well as a volatility spillover effect from the Japanese stock market to the Korean stock market.

Choudhry (2004) evaluates mean and volatility spillovers between the stock exchanges of political rivals (Israel and Jordan, India and Pakistan, Greece and Turkey) and allies (the US). The results indicate bidirectional mean and volatility spillovers between two rivals and their ally (the US), but only in the short run. Dedi and Yavas (2016) examine the linkages between equity market return and volatility spillovers and the corresponding country's market (Germany, the UK, China, Russia and Turkey). The results indicate significant co-movement of returns among the sample, with Russia and Turkey exhibiting the highest volatility, while the UK and China have

the lowest volatility spillover. All the stock exchanges except the UK and Turkey indicate volatility spillovers from other markets. The authors also find that market volatility has a positive effect on future returns only in the case of the UK.

Ely (2015) examines mean and volatility spillovers among stock markets and foreign exchange markets in Brazil. The author shows that the stock market drives the foreign exchange market in terms of price formation, with nonlinear Granger causalities from the exchange market to the stock market. Some portion of these nonlinear causalities is explained by volatility spillovers. Exchange rate volatility affects not only stock market volatility, but also stock returns. Asian markets are more sensitive to financial crises, which have a significant effect on their stock returns.

Singhania and Anchalia (2013) explore the impact of global crises on the volatility of stock returns. They find that the sub-prime crisis had a positive impact on volatility and returns in Japan, China and India, but not Hong Kong. In addition, the Euro-zone debt crisis had a negative impact on the volatility of highly volatile stock returns, such as in India and China, but this did not hold for Japan and Hong Kong. Finally, Valadkhani and Chen (2014) examine the dynamic and switching effects of volatility spillovers from the US stock market to the Australian, Canadian and UK stock markets. The study finds that US stock market volatility influences that of the other stock markets and has a significant impact on their output volatilities.

The literature review shows that the US and UK markets significantly influence regional, local and emerging markets due to financial convergence. There is also an interrelationship between return and volatility spillovers in these markets and in the PSX. We argue that CPEC and China's acquisition of 40 percent of PSX shareholdings has led to a paradigm shift, whereby changes in return and market volatility in China could now affect return and market volatility in Pakistan.

3. Econometric Methodology

We use the two-stage GARCH-in-mean approach (GARCH-M) to measure the spillover effects of international and regional markets on the PSX. Earlier studies have used different techniques to explain volatility spillovers between financial markets, including the GARCH, VAR-GARCH, M-GARCH, E-GARCH and GARCH-BEKK models.[2] We employ the

[2] See Alikhanov (2013), Carsamer (2016), Choudhry (2004), Peng et al. (2017), Dedi and Yavas (2016) and Joshi (2011).

GARCH-M model specified by Pan (1997) because it provides information on volatility spillovers without having to select explanatory variables at the outset. The moving window utility allows us to proceed without specifying any break-point, which other methodologies generally require. The strength and direction of the relationship also makes this technique a far better one.

The study examines the spillover effect of international and regional markets on the PSX, using both event periods. For regional connectivity, we consider the Shenzhen Stock Exchange (SZSE), a leading Chinese index.[3] Both these indices have been used elsewhere to examine the effect of investment. Drawing on daily stock market index data for the period 2001–17, we divide the convergence and co-integration of stock markets into three phases: (i) 2001–17 (overall sample), (ii) pre-FTA and post-FTA, and (iii) pre-CPEC and post-CPEC. We use daily rather than monthly data because the latter provides no evidence of short-run integration, while using daily data yields strong evidence of mean and volatility spillovers.[4]

To determine the mean and volatility spillover effects of international markets on the PSX and whether it follows international or regional stock exchanges, we measure the returns of the sample stock exchanges as follows:[5]

$$r_t = log\left(\frac{P_1}{P_0}\right) \tag{1}$$

where r_t is the average market return of each stock exchange for day t, and P_1 and P_0 represent the overall weighted index. To gauge the behavior and volatility of these international indices and PSX returns, the series is modeled using the following ARMA (1, 1)-GARCH (1, 1)-M model:

$$r_{i,t} = \beta_0 + \beta_1 r_{i,t-1} + \beta_2 V_{i,t} + \beta_3 \mu_{i,t-1} + \mu_{i,t} \tag{2}$$

where $\mu_{i,t} \sim N\left(0, V_{i,t}\right)$

$$V_{i,t} = \eta_0 + \eta_1 V_{i,t-1} + \eta_2 \mu_{i,t-1}^2 \tag{3}$$

where $r_{i,t}$ is the daily return of stock index k (FTSE, Nasdaq and SZSE) at time t, and $\mu_{i,t}$ is the residual, relying on the assumption that it is normally distributed with mean 0 and time-conditional variance $V_{i,t}$. We include an

[3] We recognize that not using the Shanghai Stock Exchange is a limitation of this study.
[4] The decision to use daily data is also a limitation of this study.
[5] The daily market return for each index is calculated as $r_t = \log\left(\frac{index_t}{index_{t-1}}\right)$.

ARMA (1, 1) and/or MA (1) structure in the model to adjust for possible serial correlation in the data. To measure the spillover effects of mean returns and volatility across market shares, we calculate the standardized residual and its square, and include these in the market returns and volatility equation as given below:

$$r_{j,t} = \alpha_0 + \alpha_1 r_{j,t-1} + \alpha_2 V_{j,t} + \alpha_3 \varepsilon_{j,t-1} + \varphi_j \varepsilon_i + \varepsilon_{i,t} \quad (4)$$
$$\text{where } \varepsilon_{j,t} \sim N\left(0, V_{j,t}\right)$$

$$V_{j,t} = \delta_0 + \delta_1 V_{j,t-1} + \delta_2 \varepsilon_{j,t-1}^2 + \lambda_j \varepsilon_{i,t}^2 \quad (5)$$

Here, $\varepsilon_{i,t}$ is the standardized residual series for different international indices and extracts the mean return spillover effect from these indices. To measure the spillover effects of volatility, the exogenous variable $\varepsilon_{i,t}^2$, the square of the standardized residual series (ut^2), is included in the conditional volatility equation.

Next, we examine the co-movement of stock exchanges using the Granger causality test and Johansen co-integration test. The first step is to determine whether the time-series variables are nonstationary with unit root I(1) and I(0), for which purpose we use the augmented Dickey–Fuller (ADF) and Phillips–Perron (PP) tests. In Granger's (1986) model, if the values of X predict the value of Y in a confined way, it implies that x causes y. The variable fails to Granger-cause y_t if

$$Pr(y_{t+m}|\delta) = Pr(y_{t+m}|\gamma_t) \quad (6)$$

where $Pr(y_{t+m}|\delta)$ identifies the conditional probability of y_t while x_t is the set of all information available at time t. $Pr(y_{t+m}|\gamma_t)$ indicates the conditional probability of y_t obtained by excluding all information on x_t from y_t. This set of information is represented by W_t. The causal relationship between the stationery series is estimated as:

$$x_t = \beta_0 + \sum_{j=1}^{k} \delta_j \, x_{t-j} + \sum_{j=1}^{k} \eta_j y_{t-j} + \mu_{xt} \quad (7)$$

$$y_t = \beta_0 + \sum_{j=1}^{k} \delta_j \, x_{t-j} + \sum_{j=1}^{k} \eta_j y_{t-j} + \mu_{yt} \quad (8)$$

where k is a positive integer, δ_j and η_j (j = 0, 1, ..., k) are parameters, $\beta$is a constant and μ_t is a disturbance term with zero means and finite variances. The null hypothesis states that y does not cause x. The lag is gauged using the Akaike information criterion (AIC) to measure co-movement by defining a vector of n potentially endogenous variables, Z_t.

$$Z = AiZ_{a-1} \ldots + A_k Z_{t-k} + \delta D_t + \mu_{et} \quad (9)$$

where A_i is the coefficients' matrix, n is a constant, D_t denotes seasonal dummies orthogonal to the constant term and μ is assumed to be an independent and identically distributed Gaussian process.

To determine how the sample stock exchanges reacted to the financial crisis of 2007/08, we use the seasonality method, which gauges the behavior and movement of time-series data that has experienced significant changes in a specific period in an extraordinary situation. In this case, seasonality refers to fluctuations in stock indices that occurred due to the financial crisis of 2007/08 (Fedderke & Marinkov, 2016). We use OLS to determine index behavior during the crisis. The dependent variable is the stock index. The independent variables are divided into three categories: pre-crisis (BC1–BC5), crisis (2007/08) (CRS1–CRS2) and post-crisis (2009–13) (ACRS1–ACRS5). These are quantified as 1 for the same year and 0 otherwise (Hussain et al., 2015). The econometric equation used is:

$$Y_t = \varphi_0 + \varphi_1 D_{t\ Bc1-Bc6} + \varphi_2 D_{t\ CRS1-CRS2} + \varphi_3 D_{t\ ACRS1-ACRS9} + \mu_t \quad (10)$$

The model includes only qualitative regressor-different categories of years, assigning a value of 1 if the observation belongs to that year and 0 otherwise. Y_t is the average annual index, and D_t = 1 if the observation relates to the same year and 0 otherwise.

We apply the F-test to establish the pre-sample and post-sample variance in each data stream. The t-test is used to compare statistical models as per the dataset provided.

$$F\ value = \frac{\sigma_1^2}{\sigma_2^2} \quad (11)$$

To apply an F distribution under the null hypothesis, we determine the mean of two given observations (pre-event and post-event) and then calculate their variance:

$$\sigma^2 = \frac{\sum(x-\bar{x})^2}{n-1} \quad (11.1)$$

where σ^2 is the variance, x denotes the values given in a set of data, $\bar{x}$ is the mean data and n is the total number of observations.

4. Findings and Discussion

Table 1 provides descriptive statistics for the entire sample. The PSX trends upward, growing continuously. The SZSE, a developing stock market, shows extraordinary performance.

Table 1: Descriptive statistics for stock indices, 2001–17

	SZSE	PSX	FTSE 100	Nasdaq
Mean	827	15,082	2,757	5,588
Maximum	3,141	50,192	5,914	7,103
Minimum	237	1,075	1,114	3,287
SD	431	12,151	1,074	872
Skewness	1.06	1.04	0.87	-0.45
Kurtosis	5.62	3.06	2.65	2.18
Jarque–Bera	1,958.40	740.174	538.931	252.618
Prob.	0.000	0.000	0.000	0.000

Source: Authors' calculations

4.1. Overall Sample, 2001–17

We use the GARCH variance equation to test the mean return spillover and volatility of each index. Table 2 gives the results for the overall sample, indicating that the PSX captures the mean and volatility effects of the FTSE 100 (mean (φ) = –0.0087, volatility (λ) = 0.0005 at $p < 0.05$) and Nasdaq (mean (φ) = 0.0074, volatility (λ) = 0.0003) at $p < 0.05$). We can infer, therefore, that the PSX does not capture the mean and volatility effects of the SZSE (mean (φ) = –0.0080, volatility (λ) = 0.00003 at $p > 0.05$).

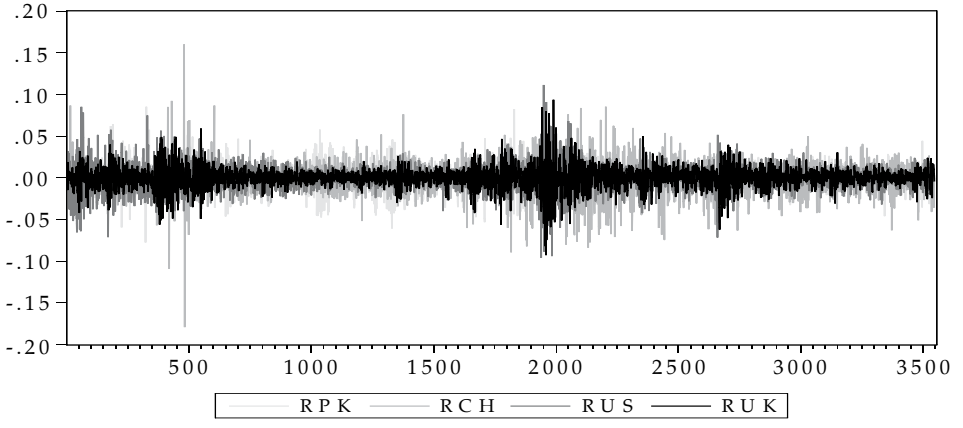
Table 2: Mean and volatility spillovers from global and regional markets to PSX, 2001–17

	FTSE 100 vs PSX		Nasdaq vs PSX		SZSE vs PSX	
	FTSE 100	**PSX**	**Nasdaq**	**PSX**	**SZSE**	**PSX**
β_0	-0.1804	0.0011	-0.0009	0.0011	0.0002	0.0011
	(2.1605)	(0.0002)	(0.0007)	(0.0002)	(0.0001)	(0.0002)
β_1	-0.0009	0.1746	2.4472	0.1746	-0.0479	0.1746
	(0.0001)	(0.1702)	(1.0619)	(0.1702)	(0.2770)	(0.1702)
β_2	2.4215	-0.0622	-2.4639	-0.0622	0.1079	-0.0622
	(1.0792)	(0.1728)	(1.0634)	(0.1728)	(0.2758)	(0.1728)
β_3	-2.4382		-0.0009		-0.0044	
	(1.0812)		(0.0008)		(1.006)	
φ		-0.0087**		0.0074*		-0.0080
		(0.0104)		(0.0113)		(0.0095)
δ_0	0.0001	0.0040	0.0001	0.0004	-0.0001	0.0004
	(0.0024)	(0.0055)	(0.0002)	(0.0005)	(0.0001)	(0.0001)
δ_1	0.0708**	0.1851**	0.0709**	0.1851**	0.1334**	0.1851**
	(0.0100)	(0.0099)	(0.0100)	(0.0099)	(0.0023)	(0.0099)
δ_2	0.8487**	0.7689**	0.8486**	0.7689**	0.8882**	0.7689**
	(0.0200)	(0.0087)	(0.0200)	(0.0087)	(0.0014)	(0.0087)
λ		0.0005**		0.0003**		0.0001**
		(0.0003)		(0.0002)		(0.0001)
			Residual diagnostics			
Q (24)	0.531	0.129	0.331	0.129	0.014	0.129
Q (24) sq.	0.976	0.189	1.000	0.189	0.791	0.189

Note: This table gives the daily market returns of the PSX, Nasdaq, FTSE 100 and SZSE from January 2001 to January 2017, using an ARMA (1, 1)-GARCH (1, 1)-M model. Standard errors are reported in parentheses. The residual diagnostics are presented as p-values. Q (24) is the portmanteau statistic, which explains the null hypothesis of no residual serial correlations, gauged through a lag of 24. Using the same method, Q (24) sq. indicates the null hypothesis of no residual ARCH effect. * and ** denote significance at 95 percent ($p < 0.05$) and 99 percent ($p < 0.01$), respectively.
Source: Authors' calculations.

These findings are in line with convergence theory, suggesting that small local markets such as the PSX converge on their financial origin, that is, industrialized and digitalized markets such as the Nasdaq and FTSE 100. Since the integration of these markets cannot occur in isolation, any systematic and unsystematic risks in the global market will not only affect returns, but also create volatility in regional and local markets (Rajhans & Jain, 2015). This finding implies that Pakistan's economy was integrated financially and socially with the US and UK markets during 2001–17. Any substantial changes in these markets had a significant impact on the PSX (Figure 1).

Figure 1: Volatility of market returns, 2001–17



We apply the ADF and PP tests to measure co-movement among the stock exchange indices. The results for all the indices over the period 2001–17 are nonstationary at level and stationary at first difference (Table A1 in the Appendix). To measure the lag length, we use the log likelihood method, AIC an Schwarz information criterion (SIC). The results indicate an appropriate lag order of 3 and the vector decomposition is thus estimated using this lag order (Table A2 in the Appendix).

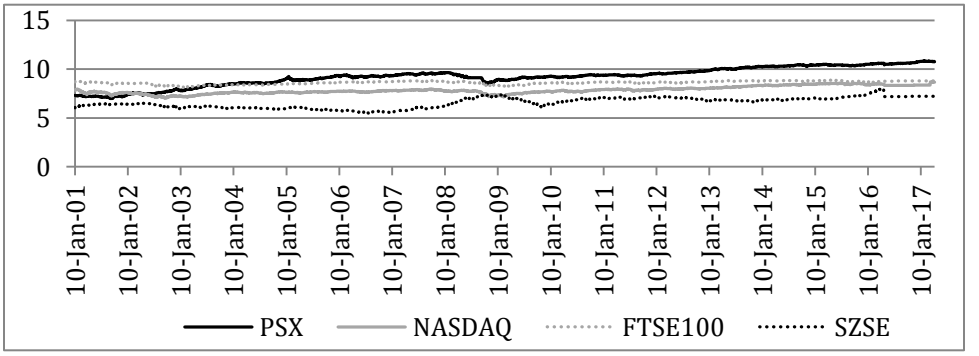
Table 3 presents the short-run co-movements of the PSX with the SZSE, Nasdaq and FTSE 100 (0.0016, $p < 0.05$; 0.0005, $p < 0.05$; and –0.0081, $p < 0.05$, respectively) as well as its long-run co-movements with the SZSE, Nasdaq and FTSE 100 (0.0079, $p < 0.001$; 0.0011, $p < 0.001$; and 0.0023, $p < 0.001$, respectively). The results imply that the co-movements in this series are negative in the short term and medium term, but positive and persistent in the long term. The results for the mean and volatility spillovers and co-movement series support convergence theory, demonstrating that small domestic markets such as the PSX converge on their financial origin, that is, industrialized, digitalized markets such as the Nasdaq and FTSE 100 (Figure 2).

Table 3: Co-movement of indices, 2001–17

	PSX	Nasdaq	FTSE 100	SZSE		PSX	Nasdaq	FTSE 100	SZSE
PSX (-1)	1.1300	-0.0016	0.0005	-0.0081	FTSE100 (-1)	0.0665	0.0060	0.9439	0.0133
	-0.0167	-0.0037	-0.0052	-0.0020		-0.0501	-0.0110	-0.0156	-0.0059
	[67.871]	[-0.436]	[0.098]	[-4.100]		[1.327]	[0.544]	[60.37]	[2.253]
PSX (-2)	-0.1621	-0.0054	-0.0017	0.0057	FTSE100 (-1)	-0.0856	0.0060	0.0109	-0.0110
	-0.0253	-0.0056	-0.0079	-0.0030		-0.0690	-0.0152	-0.0215	-0.0082
	[-6.415]	[-0.971]	[-0.212]	[1.911]		[-1.240]	[0.394]	[0.506]	[-1.347]
PSX (-3)	0.0330	0.0079	0.0011	0.0023	FTSE100 (-1)	0.0147	-0.0125	0.0357	-0.0024
	-0.0167	-0.0037	-0.0052	-0.0020		-0.0500	-0.0110	-0.0156	-0.0059
	[1.968]	[2.161]	[0.217]	[1.182]		[0.294]	[-1.141]	[2.287]	[-0.399]
Nasdaq (-1)	-0.0427	0.9823	0.0092	-0.0120	SZSE (-1)	-0.1928	0.0355	-0.0501	1.1598
	-0.0720	-0.0158	-0.0225	-0.0085		-0.1419	-0.0312	-0.0443	-0.0168
	[-0.593]	[62.116]	[0.409]	[-1.414]		[-1.358]	[1.138]	[-1.130]	[69.13]
Nasdaq (-2)	-0.0809	-0.0201	0.0093	0.0028	SZSE (-2)	0.6253	0.0538	0.0010	-0.0994
	-0.1011	-0.0222	-0.0315	-0.0120		-0.2170	-0.0477	-0.0677	-0.0257
	[-0.800]	[-0.903]	[0.293]	[0.231]		[2.882]	[1.129]	[0.014]	[-3.874]
Nasdaq (-3)	0.1260	0.0293	-0.0094	0.0106	SZSE (-3)	-0.4485	-0.0936	0.0407	-0.0635
	-0.0720	-0.0158	-0.0225	-0.0085		-0.1419	-0.0312	-0.0443	-0.0168
	[1.750]	[1.852]	[-0.416]	[1.239]		[-3.161]	[-3.005]	[0.919]	[-3.787]
C	27.5393	16.8009	35.5885	-0.2417					
	-27.9858	-6.1466	-8.7331	-3.3089					
	[0.984]	[2.733]	[4.075]	[-0.073]					

Source: Authors' calculations.

Figure 2: Co-movement of market returns, 2001–17



4.2. Pre-FTA and Post-FTA

For the FTA signed between Pakistan and China in 2006, we divide the sample into pre-event and post-event subsamples, both spanning five years. The overall results show that the PSX captures the mean and volatility effects of all the other indices before and after the event. Table 4 gives the results of the GARCH variance analysis for all series before and after the event: PSX (0.578, $p < 0.001$; 0.265, $p < 0.001$); SZSE (0.087, $p < 0.001$; 0.0436, $p < 0.001$); Nasdaq (0.077, $p < 0.001$; 0.044, $p < 0.001$); and FTSE 100 (0.129, $p <$

0.001; 0.113, $p < 0.001$). The results imply that these markets are volatile and responsive to economic shocks. They also point to the level of persistency in information and its effects on volatility.

Table 4: Mean and volatility spillovers from global and regional indices to PSX, 2001–11

	Five years' data pre-FTA (Nov 2001–Nov 2006)						Five years' data post-FTA (Nov 2006–Nov 2011)					
	FTSE 100 vs PSX		Nasdaq vs PSX		SZSE vs PSX		FTSE 100 vs PSX		Nasdaq vs PSX		SZSE vs PSX	
β_0	-0.067	-0.002	-0.020	0.007	-0.003	0.001	-0.025	0.001	-0.001	0.001	-0.015	0.001
	(0.001)	(0.002)	(0.056)	(0.010)	(0.001)	(0.001)	(0.061)	(0.005)	(0.001)	(0.005)	(0.066)	(0.005)
β_1	-0.321	0.739	0.479	0.682	0.821	0.464	0.622	1.603	-1.381	1.469	2.415	1.461
	(0.190)	(0.169)	(0.473)	(0.154)	(0.681)	(0.293)	(0.772)	(1.901)	(1.610)	(1.906)	(2.151)	(1.929)
β_2	0.123	-0.522	-1.102	-0.377	-0.517	0.066	-3.094	0.029	1.328	0.525	-3.371	0.395
	(0.344)	(0.383)	(0.930)	(0.356)	(1.318)	(0.754)	(1.363)	(2.320)	(2.077)	(2.265)	(2.428)	(2.296)
β_3	-0.054	-0.009	0.518	-0.098	-0.229	-0.253	2.407	-1.601	0.091	-1.966	0.939	-1.833
	(0.021)	(0.398)	(0.782)	(0.377)	(0.978)	(0.722)	(1.151)	(1.145)	(1.050)	(1.125)	(1.025)	(1.130)
φ		0.002		-0.004		0.016		-0.084		0.082		0.045**
		(0.020)		(0.007)		(0.072)		(0.075)		(0.065)		(0.052)
δ_0	0.023	0.002	0.003	0.003	0.010	0.002	0.0027	0.002	0.004	0.001	0.004	0.001
	(0.011)	(0.002)	(0.001)	(0.002)	(0.004)	(0.008)	(0.0013)	(0.004)	(0.001)	(0.005)	(0.003)	(0.004)
δ_1	0.134	0.578	0.077	0.678	0.087	0.139	0.1133	0.265	0.044	0.272	0.0436	0.272
	(0.039)	(0.064)	(0.020)	(0.054)	(0.015)	(0.036)	(0.0382)	(0.060)	(0.024)	(0.062)	(0.032)	(0.063)
δ_2	0.456	0.593*	0.902	0.602*	0.881	0.559	0.8226	0.699*	0.901	0.693*	0.6617	0.691
	(0.012)	(0.028)	(0.020)	(0.019)	(0.023)	(0.195)	(0.0567)	(0.048)	(0.046)	(0.051)	(0.272)	(0.052)
λ		-0.045**		-0.001**		0.002**		0.078**		0.145**		0.0123**
		(0.0123)		(0.003)		(0.001)		(0.004)		(0.000)		(0.002)
Residual diagnostics												
Q (24)	0.890	0.611	0.891	0.450	0.771	0.063	0.270	0.630	0.868	0.548	0.310	0.534
Q (24) sq.	0.788	0.869	0.901	0.338	0.538	0.012	0.340	0.854	0.591	0.982	0.465	0.983

Note: Standard errors are reported in parentheses. The residual diagnostics are presented as p-values. Q (24) is the portmanteau statistic, which explains the null hypothesis of no residual serial correlations, gauged through a lag of 24. Using the same method, Q (24) sq. indicates the null hypothesis of no residual ARCH effect. * and ** denote significance at 95 percent ($p < 0.05$) and 99 percent ($p < 0.01$), respectively.
Source: Authors' calculations.

The sum of the GARCH coefficients for all the indices is close to 1 just before the pre-FTA and post-FTA points, indicating that shocks are persistent and that forecast conditional volatility takes longer to stabilize around the variance. The results suggest that all the indices exert some influence over the PSX in terms of mean spillover pre-FTA and post-FTA, but that this impact is highest for the SZSE post-FTA. The PSX also captures volatility effects pre-FTA and post-FTA: its post-FTA volatility is affected the most by the FTSE 100 and Nasdaq. Overall, the PSX appears to capture the volatility effects of these global markets during the period 2001–11 (Figures 3 and 4).

Figure 3: Pre-FTA volatility of returns

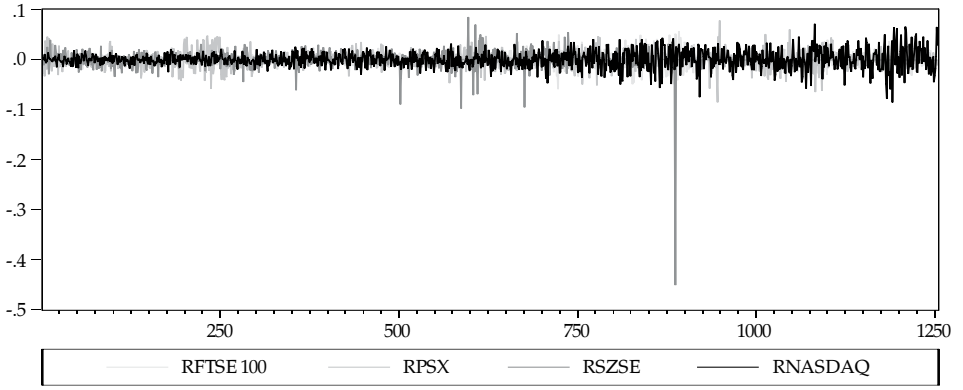
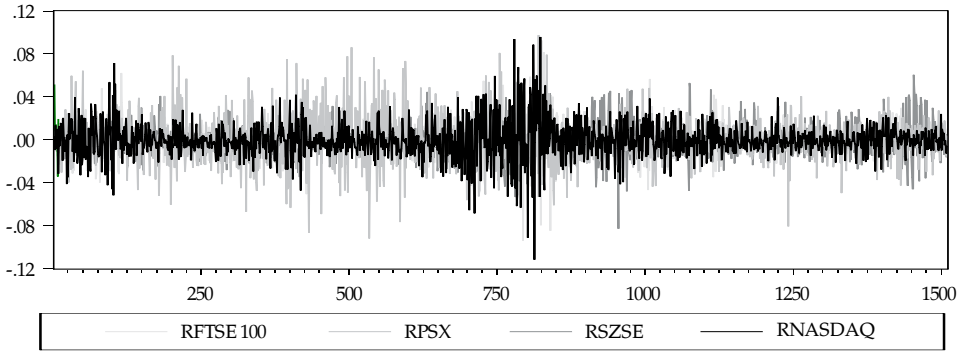


Figure 4: Post-FTA volatility of returns



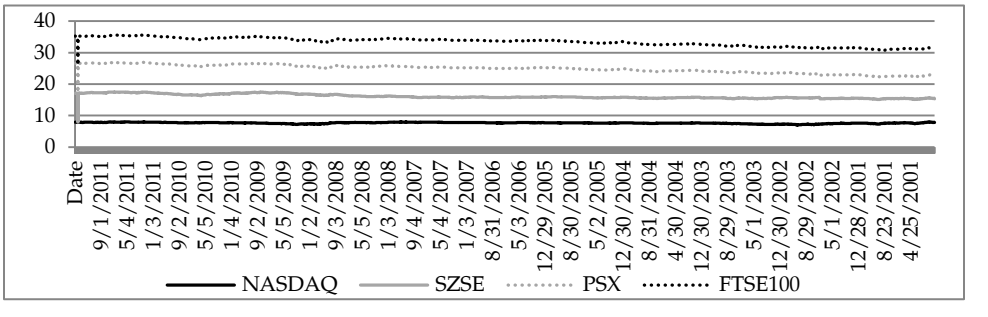
We use the ADF and PP tests to measure the co-movement among the indices. The results for all the indices over the period 2013–17 are nonstationary at level and stationary at first difference (Table A3 in the Appendix). To measure the lag length, we use the log likelihood, AIC and SIC (Table A4 in the Appendix). The results in Table 5 indicate the short-run and long-run co-movements of the PSX with all the indices under observation during 2001–11. While the co-movements in the short term and medium term are negative, they are positive and persistent in the long run (Figure 5). The results for mean and volatility spillovers and co-movements in the series support convergence theory, implying that small domestic markets such as the PSX converge on their financial origin, that is, industrialized, digitalized markets such as the Nasdaq and FTSE 100.

Table 5: Co-movement of indices, 2001–11

	PSX	SZSE	Nasdaq	FTSE 100		PSX	SZSE	Nasdaq	FTSE 100
PSX (-1)	1.130	-0.008	-0.002	0.001	SZSE (-1)	-0.193	1.160	0.035	-0.050
	-0.017	-0.002	-0.004	-0.005		-0.142	-0.017	-0.031	-0.044
	[67.874]	[-4.100]	[-0.436]	[0.098]		[-1.358]	[69.131]	[1.138]	[-1.130]
PSX (-2)	-0.162	0.006	-0.005	-0.002	SZSE (-2)	0.625	-0.099	0.054	0.001
	-0.025	-0.003	-0.006	-0.008		-0.217	-0.026	-0.048	-0.068
	[-6.415]	[1.911]	[-0.971]	[-0.212]		[2.882]	[-3.874]	[1.129]	[0.014]
PSX (-3)	0.033	0.002	0.008	0.001	SZSE (-3)	-0.448	-0.064	-0.094	0.041
	-0.017	-0.002	-0.004	-0.005		-0.142	-0.017	-0.031	-0.044
	[1.968]	[1.182]	[2.161]	[0.217]		[-3.161]	[-3.787]	[-3.005]	[0.919]
Nasdaq (-1)	-0.043	-0.012	0.982	0.009	FTSE100 (-1)	0.0665	0.0133	0.0060	0.9439
	-0.072	-0.009	-0.016	-0.022		-0.0501	-0.0059	-0.0110	-0.0156
	[-0.593]	[-1.414]	[62.116]	[0.409]		[1.327]	[2.253]	[0.544]	[60.372]
Nasdaq (-2)	-0.081	0.003	-0.020	0.009	FTSE100 (-2)	-0.0856	-0.0110	0.0060	0.0109
	-0.101	-0.012	-0.022	-0.032		-0.0690	-0.0082	-0.0152	-0.0215
	[-0.800]	[0.231]	[-0.907]	[0.293]		[-1.240]	[-1.347]	[0.394]	[0.506]
Nasdaq (-3)	0.126	0.011	0.029	-0.009	FTSE100 (-3)	0.0147	-0.0024	-0.0125	0.0357
	-0.072	-0.009	-0.016	-0.022		-0.0500	-0.0059	-0.0110	-0.0156
	[1.750]	[1.239]	[1.852]	[-0.416]		[0.294]	[-0.399]	[-1.141]	[2.287]
C	27.539	-0.242	16.801	35.589					
	-27.986	-3.309	-6.147	-8.733					
	[0.984]	[-0.073]	[2.733]	[4.075]					

Note: The figures below the estimates are p-values and those in brackets are standard errors. The lag length of co-integration is set at 3. The trace statistics indicate two co-integrating equations at the 0.05 significance level.
Source: Authors' calculations.

Figure 5: Co-movement of market returns, 2001–11



4.3. Pre-CPEC and Post-CPEC

The CPEC agreement between Pakistan and China was signed on 20th April 2015 and aims to bolster trade between the two countries. Given its likely influence over the PSX, we investigate the latter's convergence on international or regional stock markets. Table 6 gives the results of the GARCH mean and variance analysis for all pre-CPEC and post-CPEC series. The sum of the GARCH coefficients of all the indices is almost 1 in both cases (pre-event and post-event), indicating that shocks are persistent and that

forecast conditional volatility takes longer to stabilize around the variance (Figures 6 and 7). Both series suggest that the PSX does not capture the pre-event or post-event mean effects of the indices, barring the SZSE. This implies that, post-CPEC, the PSX has shifted toward the regional market.

Table 6: Mean and volatility spillovers from global and regional indices to PSX, 2013–17

	Two years' data post-CPEC (Apr 2015–Apr 2017)						Two years' data pre-CPEC (Apr 2013–Apr 2015)					
	FTSE 100 vs PSX		Nasdaq vs PSX		SZSE vs PSX		FTSE 100 vs PSX		Nasdaq vs PSX		SZSE vs PSX	
β_0	-0.004	-0.001	-0.003	-0.006	-0.002	-0.006	0.004	-0.001	-0.003	-0.001	0.004	-0.001
	(0.001)	(0.002)	(0.002)	(0.002)	(0.001)	(0.002)	(0.011)	(0.002)	(0.004)	(0.002)	(0.002)	(0.002)
β_1	0.485	3.432	-3.460	3.403	-0.494	3.399	17.226	0.002	-2.021	-0.038	-8.166	-0.014
	(0.136)	(8.007)	(3.217)	(8.332)	(1.050)	(8.496)	(41.607)	(0.997)	(4.160)	(0.996)	(5.043)	(0.995)
β_2	-1.918	-2.886	2.553	-2.858	1.504	-2.854	-28.257	0.999	4.163	1.109	7.907	1.023
	(0.253)	(7.777)	(3.461)	(8.077)	(2.244)	(8.243)	(42.411)	(4.493)	(4.491)	(4.505)	(5.027)	(4.487)
β_3	1.286	-0.438	0.890	-0.442	-0.965	-0.443	11.031	-0.882	-2.131	-0.961	0.269	-0.891
	(0.212)	(0.847)	(0.938)	(0.859)	(2.015)	(0.856)	(4.904)	(4.413)	(1.690)	(4.429)	(0.661)	(4.419)
φ		0.045		0.002		-0.012*		-0.025		0.090		0.011
		(0.198)		(0.141)		(0.071)		(0.183)		(0.136)		(0.084)
δ_0	0.001	0.010	0.004	0.010	0.004	0.010	0.005	0.008	0.002	0.008	0.014	0.008
	(0.007)	(0.006)	(0.001)	(0.006)	(0.002)	(0.006)	(0.002)	(0.009)	(0.010)	(0.009)	(0.011)	(0.009)
δ_1	-0.006	0.150	0.139	0.150	0.083	0.150	0.163	0.150	0.116	0.150	0.019	0.150
	(0.002)	(0.169)	(0.029)	(0.175)	(0.022)	(0.175)	(0.037)	(0.144)	(0.041)	(0.145)	(0.018)	(0.146)
δ_2	1.047	0.600	0.826	0.600	0.909	0.600	0.759	0.600	0.756	0.600	0.909	0.600
	(0.007)	(0.268)	(0.032)	(0.267)	(0.023)	(0.271)	(0.054)	(0.392)	(0.089)	(0.393)	(0.102)	(0.394)
λ		0.009		0.008		0.023**		0.002*		0.003*		0.002
		(0.006)		(0.008)		(0.011)		(0.001)		(0.001)		(0.001)
Residual diagnostics												
Q (24)	0.124	0.704	0.358	0.740	0.973	0.712	0.596	0.476	0.854	0.370	0.843	0.443
Q (24) sq.	0.108	1.000	0.872	1.000	0.856	0.514	0.925	0.019	0.859	0.026	0.956	0.018

Note: Standard errors are reported in parentheses. The residual diagnostics are presented as p-values. Q (24) is the portmanteau statistic, which explains the null hypothesis of no residual serial correlations, gauged through a lag of 24. Using the same method, Q (24) sq. indicates the null hypothesis of no residual ARCH effect. * and ** denote significance at 95 percent ($p < 0.05$) and 99 percent ($p < 0.01$), respectively.
Source: Authors' calculations.

Figure 6: Pre-CPEC volatility of returns

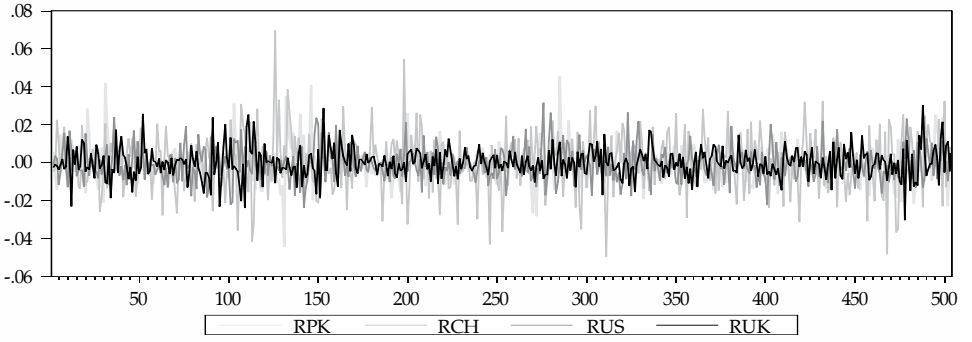
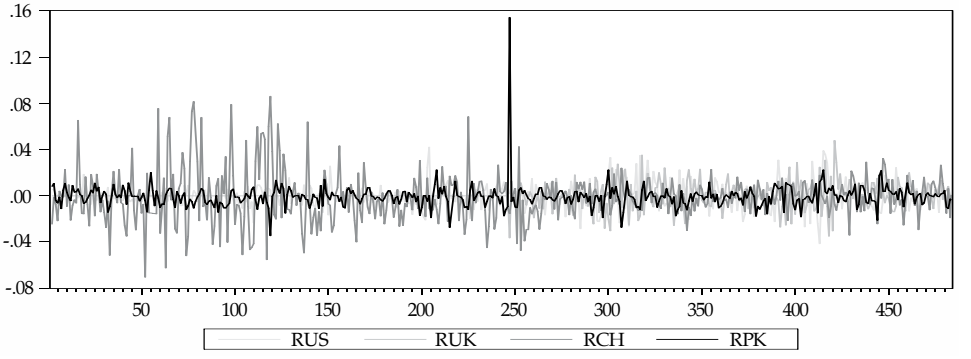


Figure 7: Post-CPEC volatility of returns



We use the ADF and PP tests to measure the co-movement among the indices. The results show that all the indices are nonstationary at level and stationary at first difference over the period 2013–17 (Table A5 in the Appendix). To measure the lag length, we use the log likelihood, AIC and SIC (Table A6 in the Appendix). Table 7 indicates the co-movement of the PSX with all the indices. The results suggest that the PSX follows these indices, although the mean, volatility and co-movement results for the post-CPEC period point to a paradigm shift (Figure 8), that is, the PSX tends to converge on the SZSE post-CPEC.

Table 7: Co-movement of indices, 2013–17

	PSX	Nasdaq	FTSE 100	SZSE		PSX	Nasdaq	FTSE 100	SZSE
PSX (-1)	1.1191	-0.0108	0.0069	-0.0246	FTSE100 (-1)	0.5296	0.0848	0.9583	0.0406
	-0.0509	-0.0114	-0.0076	-0.0067		-0.2821	-0.0629	-0.0422	-0.0372
	[21.971]	[-0.952]	[0.902]	[-3.653]		[1.877]	[1.347]	[22.702]	[1.089]
PSX (-2)	-0.1214	0.0116	-0.0065	0.0247	FTSE100 (-2)	-0.5810	-0.0811	0.0179	-0.0622
	-0.0509	-0.0114	-0.0076	-0.0067		-0.2835	-0.0633	-0.0424	-0.0374
	[-2.382]	[1.020]	[-0.853]	[3.666]		[-2.049]	[-1.281]	[0.422]	[-1.661]
Nasdaq (-1)	-0.1219	0.9907	-0.0141	-0.0346	SZSE (-1)	0.0549	0.1341	-0.0552	1.2984
	-0.1949	-0.0435	-0.0292	-0.0258		-0.3848	-0.0859	-0.0575	-0.0508
	[-0.625]	[22.771]	[-0.484]	[-1.343]		[0.142]	[1.562]	[-0.959]	[25.541]
Nasdaq (-2)	0.0504	-0.0017	0.0173	0.0356	SZSE (-2)	-0.1101	-0.1420	0.0492	-0.3170
	-0.1971	-0.0440	-0.0295	-0.0260		-0.3828	-0.0854	-0.0572	-0.0505
	[0.255]	[-0.038]	[0.587]	[1.368]		[-0.287]	[-1.661]	[0.859]	[-6.268]
C	853.0373	10.6367	134.1324	159.7401					
	-493.3200	-110.1200	-73.8021	-65.1662					
	[1.729]	[0.096]	[1.817]	[2.451]					

Source: Authors' calculations.

Figure 8: Co-movement of market returns, 2013–17

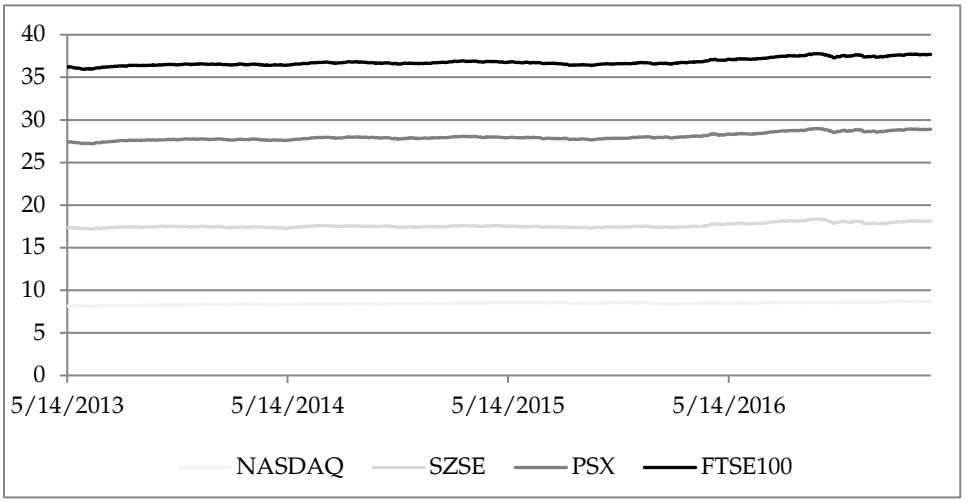


Table 8 gives the results of the Granger causality test, indicating that the PSX followed the SZSE, FTSE 100 and Nasdaq from 2006 to 2017.

Table 8: Granger causality test results

	Post-FTA		Post-CPEC	
	F-stat	Prob.	F-stat	Prob.
SZSE does not Granger-cause PSX	3.7945	0.0099	0.4122	0.6623
PSX does not Granger-cause SZSE	8.2670	2.00E-05	7.5076	0.0006
Nasdaq does not Granger-cause PSX	0.6782	0.5653	2.4582	0.0865
PSX does not Granger-cause Nasdaq	9.1497	5.00E-06	0.5424	0.5816
FTSE 100 does not Granger-cause PSX	0.7664	0.5127	2.0398	0.1310
PSX does not Granger-cause FTSE 100	2.9646	0.0309	0.1721	0.8419

Source: Authors' calculations.

4.4. Financial Crisis

Financial crises will, naturally, affect stock market activities. This study considers the financial crisis of 2007/08, using OLS and a qualitative regressor. The dependent variable is the stock index and the independent variables are divided into three categories: pre-crisis period (BC1–BC5), crisis period (CRS1–CRS2) and post-crisis period (ACRS1–ACRS5). Table 9 shows that all the stock indices trended downward during the financial crisis period, which implies that the crisis affected their stability, soundness and responsiveness to economic shocks. The Nasdaq and PSX appear to have been affected less than the FTSE 100 and SZSE (Figure 9). The key factor is

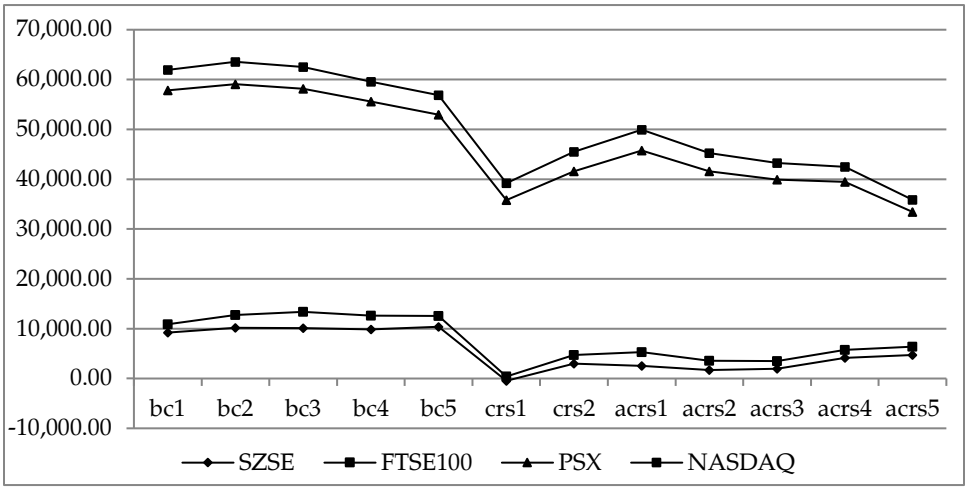
the soundness and stability of international institutional investors and integration with stable international markets.

Table 9: Crisis effects

	SZSE	FTSE 100	PSX	Nasdaq
bc1	9,186.718	1,695.930	46,954.019	4,077.659
bc2	10,169.799	2,594.241	46,304.270	4,492.606
bc3	10,098.984	3,281.297	44,789.807	4,353.138
bc4	9,856.796	2,774.718	42,943.060	4,019.510
bc5	10,384.464	2,165.949	40,421.865	3,911.774
crs1	-459.658	8,600.906	35,396.070	3,424.790
crs2	2,995.280	1,740.914	36,871.581	3,863.429
acrs1	2,509.926	2,754.980	40,515.004	4,155.847
acrs2	1,690.975	1,888.968	38,013.063	3,678.469
acrs3	1,962.153	1,511.497	36,448.124	3,311.800
acrs4	4,153.195	1,577.922	33,699.082	3,028.149
acrs5	4,702.849	1,695.922	27,030.689	2,437.235

Note: * $p < 0.05$, ** $p < 0.01$.
Source: Authors' calculations.

Figure 9: Market indices growth during financial crisis period



4.5. Pre-Event and Post-Event Variance Analysis

This section applies the F-test to examine the distribution of the variances of two populations, that is, to gauge if the pre-selection and post-selection of events is equal. Table 10 reports the results of the F-test and shows that, pre-FTA and post-FTA for each index, there is strong evidence

in favor of rejecting the null hypothesis. This implies that the pre-FTA and post-FTA variances are not equal at the 5 percent significance level. It also suggests that the increase in overall performance of the PSX is due to its connection with the Chinese stock market.

Table 10: F-test results

Event	Stock exchange	Pre-event	Post-event	F-value	(2*Pr(F < f)) Ha: ratio = 1
FTA 2006	PSX	4,315.324	7,934.807		
		(2,715.511)	(4,097.611)	0.4392	0.000
	Nasdaq	1,860.721	2,107.501		
		(308.6315)	(397.1716)	0.6038	0.000
	FTSE 100	4,761.234	5,200.969		
		(616.894)	(773.4536)	0.6361	0.000
	SZSE	3,257.996	5,901.661		
		(720.009)	(4,326.349)	0.0277	0.000
Financial crisis 2007/08	PSX	6,182.441	18,917		
		(3,238.34)	(7,937.664)	0.1664	0.000
	Nasdaq	1,906.947	3,181.476		
		(322.241)	(740.524)	0.1894	0.000
	FTSE 100	4,856.567	6,016.433		
		(726.793)	(526.740)	1.9038	0.000
	SZSE	3,437.325	10,276		
		(565.065)	(1,943.049)	0.0846	0.000
CPEC 2015	PSX	38,986.58	8,963.208		
		(5,568.159)	(702.869)	62.758	0.000
	Nasdaq	5,098.824	4,290.998		
		(361.204)	(435.970)	0.6864	0.000
	FTSE 100	6,415.03	6,684.891		
		(248.526)	(180.995)	1.8854	0.000
	SZSE	10,241.64	8,963.208		
		(2,819.828)	(702.869)	16.0952	0.000

Source: Authors' calculations.

From Pakistan's perspective, the strength of the FTA lies in the considerable growth of overall exports to China since the year it was enforced. China's exports to Pakistan have also gained momentum, climbing by US$13 billion between 2006 and 2014. This substantial increase in Chinese imports is associated with the lower price of Chinese products, trade diversions due to lower transportation costs and smaller tariffs imposed by Pakistan on Chinese products. Consequently, China is now Pakistan's top importing partner.

Likewise, the pre-crisis and post-crisis statistics for each index enable us to reject the null hypothesis that the pre-event and post-event variances are not equal at the 0.05 significance level. We also find that the volatility of stock returns increased across all the markets under study. In general, the positive return spillover effects are unidirectional, running from both the US and UK to the PSX and SZSE. In the CPEC case, the post-event and pre-event variances are not equal at the 5 percent significance level, implying that Pakistan's financial and strategic integration with China has influenced the PSX, resulting in a multilateral economic order (Epaulard & Pommeret, 2016). The dependency of the PSX on the SZSE has thus increased.

Cross-border transactions through depository certificates and cross-border listings introduce further challenges and complexities, making it necessary for institutions to be well equipped and aware of the relevant legal procedures of both countries. Integration implies that the economic shocks and high volatility experienced in one economy could put pressure on its strategic partners, as in the case of the impact of the SZSE on the PSX. These factors need to be incorporated into estimations and forecasts to enable more effective risk management.

5. Theoretical Contribution and Practical Implications

This paper makes an important contribution to the theoretical literature as well as highlighting practical implications for policymakers and managers. The key insight is that, after demutualization and the CPEC agreement, trading activities were enhances when Chinese companies floated shares on the PSX through cross-listing, cross-border transactions and depository certificates. This has increased competition in the local market and created regulatory challenges, but it has also generated demand for better control management and risk management at the micro and macro levels of market operations.

Since cross-border transactions introduce additional complexities, regulatory bodies such as the PSX and Securities and Exchange Commission of Pakistan need to be aware of the legal and regulatory procedures that govern cross-listing and cross-border transactions. Financial integration includes a degree of risk, but also potentially allows risk sharing (Epaulard & Pommeret, 2016). Economic shocks to China's economy could penetrate Pakistan's capital market. This needs to be incorporated into decision making and forecasts to ensure financial stability and market performance in the future.

In addition, the study provides insights relevant to investment decisions in terms of company cross-listing and depository certificates for individual investors. We suggest that the paradigm shift will enable firms to avail a more diversified pool of resources, but that they should maintain similar levels of local and international shareholdings to accommodate international market fluctuations. Managers should be aware that market competition will increase as financial integration with China progresses. However, an additional benefit for the PSX will be the transfer of expertise and market knowledge from Chinese companies cross-listed on the Nasdaq and FTSE 100, including best practices from a regulatory perspective, technical capacity building, improved financial forecasting and investors' protection. On the other hand, managers will need to be equipped to deal with the complexities associated with cross-border transactions and business.

The findings of this study may be of interest to regulatory bodies and policymakers, who should keep in mind that local companies will now face greater competition. The rising cost of local businesses (cost-push inflation) due to the energy crisis and high levels of taxation could affect profitability and marketability vis-à-vis the more well-established Chinese firms, thereby posing a threat to local business survival. Policymakers and regulatory bodies should focus on improving the country's stock market volume and strengthening enforcement strategies, so that cross-listed firms and local firms are valued fairly. Similarly, policymakers should be aware that smaller local investors may not be equipped to deal with international market trends and strategies. Their capacity could be enhanced through outreach and training in international business strategies.

6. Conclusion

This study examines return and volatility spillovers and co-movements between the PSX and a sample of international and regional indices. We look specifically at pre-event and post-event subsamples to account for the impact of the FTA and CPEC. We find that, during 2006–15 (post-FTA), the PSX was influenced by return and volatility spillovers from global stock indices. Integration with the global market in this case implies that any systematic and unsystematic risks in the latter will affect the returns and variances of regional and local markets. Since Pakistan's economy was integrated with the US and UK markets during 2001–17, any substantial changes would have had a significant impact on the PSX.

Alternatively, when we look at the post-CPEC impact of the SZSE as a regional market, the demutualization of the PSX and strategic partnership

with the Chinese consortium, the results for mean and volatility spillovers and co-movement indicate a paradigm shift that should be closely monitored in the coming years. Pakistan's financial and strategic integration with China has clearly influenced the PSX, resulting in a multilateral economic order (Epaulard & Pommeret, 2016). The PSX is now more dependent on the SZSE. Both scenarios support convergence theory. In the post-FTA scenario, the degree of financial integration was not strong enough to affect the PSX directly, as the stock market was not directly associated with the commodities market. In the post-CPEC scenario, when the Chinese consortium acquired 40 percent of PSX shares, this had a direct impact on the stock exchange's movements and behavior.

This study contributes to an important area of research – whether regional connectivity causes the dependency of regional markets on each other. We explore significant outperformance, including index stability and soundness, to see if this proposition is justified. Our findings support the argument, indicating that the overall market efficiency and performance of regional stock exchanges is higher than before. Further research could analyze whether these stock exchanges can enhance their market performance and growth by integrating their activities with local and regional forums, and diversifying the strategic partners involved in their governance structure.

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Appendix

Table A1: Unit root test results for overall sample

	ADF test				PP test			
	Level		First difference		Level		First difference	
	With intercept	With intercept and trend	With intercept	With intercept and trend	With intercept	With intercept and trend	With intercept	With intercept and trend
PSX	1.926	-0.305	-56.765	-56.840	2.089	-0.188	-56.606	-56.627
Prob.	1.000	0.990	0.000	0.000	1.000	1.000	0.000	0.000
Nasdaq	0.271	-2.440	-64.559	-64.604	0.407	-2.337	-64.639	-64.715
Prob.	0.980	0.360	0.000	0.000	0.980	0.410	0.000	0.000
SZSE	-2.526	-3.943	-19.668	-19.666	-2.289	-3.601	-57.643	-57.636
Prob.	0.110	0.010	0.000	0.000	0.180	0.030	0.000	0.000
FTSE 100	-2.134	-3.447	-67.141	-67.141	-1.871	-3.190	-67.688	-67.696
Prob.	0.230	0.050	0.000	0.000	0.350	0.090	0.000	0.000

Note: Critical ADF test values: –3.438, –2.864 and –2.568, significant at 1, 5 and 10 percent, with only intercept. Critical PP test values: –3.970, –3.415 and –3.129, significant at 1, 5 and 10 percent, with intercept and trend.
Source: Authors' calculations.

Table A2: Lag length results for overall sample

Test	Lag	PSX	Nasdaq	FTSE 100	SZSE
Log likelihood	2	-27,502.58	-21,274.59	-22,714.22	-18,730.38
AIC		13.3908	10.3595	11.0602	9.1211
SIC		13.4047	10.3733	11.0740	9.1349
Log likelihood	3	-27,489.76	-21,262.9	-22,705.68	-18,718.84
AIC		13.3898	10.3582	11.0607	9.1196
SIC		13.4098	10.3782	11.0807	9.1396
Log likelihood	4	-27,476.65	-21,255.75	-22,698.36	-18,701.30
AIC		13.3886	10.3592	11.0617	9.1153
SIC		13.4148	10.3854	11.0879	9.1414
Hypothesized no. of CE(s)	**Eigenvalue**	**Trace statistic**	**Critical value**	**Prob.****	
None	0.0092	74.6313	47.8561	0.00	
At most 1	0.0060	36.3170	29.7971	0.00	
At most 2	0.0023	11.7390	15.4947	0.17	
At most 3	0.0005	2.07978	3.84147	0.15	

Note: At lag 3, all the test values are at a minimum level, indicating the appropriate lag order and vector decomposition. The results of the trace test indicate two co-integrating equations at the 5 percent level.
Source: Authors' calculations.

Table A3: Unit root test results for pre-FTA and post-FTA data

	ADF test				PP test			
	Level		First difference		Level		First difference	
	With intercept	With intercept and trend	With intercept	With intercept and trend	With intercept	With intercept and trend	With intercept	With intercept and trend
PSX	1.926	-0.305	-56.765	-56.840	2.089	-0.188	-56.606	-56.627
Prob.	0.9999	0.991	0.000	0.000	0.9999	0.993	0.000	0.000
Nasdaq	0.271	-2.440	-64.559	-64.604	0.407	-2.337	-64.639	-64.715
Prob.	0.977	0.359	0.000	0.000	0.983	0.413	0.000	0.000
SZSE	-2.526	-3.943	-19.668	-19.666	-2.289	-3.601	-57.643	-57.636
Prob.	0.109	0.011	0.000	0.000	0.176	0.030	0.000	0.000
FTSE100	-2.134	-3.447	-67.141	-67.141	-1.871	-3.190	-67.688	-67.696
Prob.	0.231	0.046	0.000	0.000	0.346	0.087	0.000	0.000

Note: Critical ADF test values: –3.438, –2.864 and –2.568, significant at 1, 5 and 10 percent, with only intercept. Critical PP test values: –3.970, –3.415 and –3.129, significant at 1, 5 and 10 percent, with intercept and trend.
Source: Authors' calculations.

Table A4: Lag length results for pre-FTA and post-FTA data

Test	Lag	PSX	Nasdaq	FTSE 100	SZSE
Log likelihood	2	-27,502.580	-21,274.590	-22,714.220	-18,730.380
AIC		13.391	10.360	11.060	9.121
SIC		13.405	10.373	11.074	9.135
Log likelihood	3	-27,489.760	-21,262.900	-22,705.680	-18,718.840
AIC		13.390	10.358	11.061	9.120
SIC		13.410	10.378	11.081	9.140
Log likelihood	4	-27,476.650	-21,255.750	-22,698.360	-18,701.300
AIC		13.389	10.359	11.062	9.115
SIC		13.415	10.385	11.088	9.141
Hypothesized no. of CE(s)	**Eigenvalue**	**Trace statistic**	**Critical value**	**Prob.****	
None*	0.009	74.631	47.856	0.000	
At most 1*	0.006	36.317	29.797	0.008	
At most 2	0.002	11.739	15.495	0.170	
At most 3	0.001	2.080	3.841	0.149	

Note: At lag 3, all the test values are at a minimum level, indicating the appropriate lag order and vector decomposition. The results of the trace test indicate two co-integrating equations at the 5 percent level.
Source: Authors' calculations.

Table A5: Unit root test results for pre-CPEC and post-CPEC data

	ADF test				PP test			
	Level		First difference		Level		First difference	
	With intercept	With intercept and trend	With intercept	With intercept and trend	With intercept	With intercept and trend	With intercept	With intercept and trend
PSX	-0.389	-2.009	-21.376	-21.377	-0.137	-1.815	-21.243	-21.255
Prob.	0.908	0.594	0.000	0.000	0.943	0.697	0.000	0.000
Nasdaq	-1.557	-1.188	-23.811	-23.839	-1.592	-1.236	-23.812	-23.847
Prob.	0.504	0.911	0.000	0.000	0.486	0.901	0.000	0.000
SZSE	-1.988	-1.953	-19.929	-19.919	-2.058	-2.029	-20.233	-20.222
Prob.	0.292	0.625	0.000	0.000	0.262	0.584	0.000	0.000
FTSE 100	-2.277	-2.283	-24.881	-24.859	-2.176	-2.181	-25.093	-25.069
Prob.	0.180	0.442	0.000	0.000	0.215	0.499	0.000	0.000

Note: Critical ADF test values: –3.438, –2.864 and –2.568, significant at 1, 5 and 10 percent, with only intercept. Critical PP test values: –3.970, –3.415 and –3.129, significant at 1, 5 and 10 percent, with intercept and trend.
Source: Authors' calculations.

Table A6: Lag length results for pre-CPEC and post-CPEC data

Test	Lag	PSX	Nasdaq	FTSE 100	SZSE
Log likelihood	2	-4,243.794	-3,374.032	-3,141.927	-3,069.748
AIC		14.6648	11.6656	10.8652	10.6163
SIC		14.7325	11.7333	10.9329	10.6840
Log likelihood	3	-4,231.098	-3,363.696	-3,135.605	-3,060.667
AIC		14.6601	11.6638	10.8760	10.6171
SIC		14.7580	11.7618	10.9739	10.7150
Log likelihood	4	-4,220.123	-3,357.579	-3,125.88	-3,051.505
AIC		14.6613	11.6767	10.8750	10.6176
SIC		14.7895	11.8049	11.0032	10.7458
Hypothesized no. of CE(s)	**Eigenvalue**	**Trace statistic**	**Critical value**	**Prob.****	
None	0.0224	28.4997	47.8561	0.79	
At most 1	0.0158	15.3325	29.7970	0.75	
At most 2	0.0061	6.0612	15.4947	0.68	
At most 3	0.0042	2.4806	3.8414	0.11	

Note: At lag 2, all the test values are at a minimum level, indicating the appropriate lag order and vector decomposition. The results of the trace test indicate two co-integrating equations at the 5 percent level.
Source: Authors' calculations.

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Exchange Rate Exposure and Firm Value: An Assessment of Domestic Versus Multinational Firms

Hajra Ihsan*, Abdul Rashid and Anam Naz*****

Abstract

This paper examines the impact of exchange rate changes on the stock returns of 232 nonfinancial firms listed on the Pakistan Stock Exchange, for the period January 2000 to June 2014. To mitigate the problem of heteroskedasticity, we use a generalized least squares estimator. The estimated regression models indicate that exchange rate variations have a significant effect on firm value and that firms are exposed significantly to one-period lagged variation in the exchange rate. Our results suggest that, in addition to exchange rate dynamics, increased exchange rate volatility appears to have significant and negative effects on firms' stock returns. Compared to domestic firms, multinational firms experience greater exchange rate exposure. Finally, we show that exchange rate depreciation and appreciation have significant differential effects on firms' stock returns. These effects vary significantly across domestic and multinational firms.

Keywords: Exchange rate exposure, stock returns, firm value, domestic firms, multinational firms, volatility of exchange rate, Pakistan.

JEL classification: G32.

1. Introduction

Since the breakdown of the Bretton Woods fixed exchange rate system in 1973, the influence of exchange rate levels and volatility on firm value has become increasingly important in international financial management (see Allayannis et al., 2001; Papaioannou, 2001, 2006; Jongen et al., 2012). According to standard economic theory, exchange rate volatility affects both the market value of firms as well as cash flows from their operations. The wide currency fluctuations observed over the last few

* Assistant Professor, Department of Economics, International Institute of Islamic Economics, International Islamic University, Islamabad, Pakistan.
** Associate Professor, Department of Economics, International Institute of Islamic Economics, International Islamic University, Islamabad, Pakistan.
*** MS Scholar, International Institute of Islamic Economics, International Islamic University, Islamabad, Pakistan.

decades have reinforced the interest in empirically exploring the potential exposure of multinational and domestic firms to foreign exchange rates.

The literature examines several aspects of firms' exposure to exchange rate changes and the associated risks. Bartov and Bodnar (1994), Levi (1994) and Flota (2014) look at the exchange rate effect on firms' market value, while Baum et al. (2001) consider the impact on profitability. Bartram (2008) examines the effect on hedging activities and cash flows. Jorion (1991), Granger et al. (2000), Muhammad and Rasheed (2002), Rashid (2010), and Pan and Tang (2012) look at how exchange rate changes affect stock returns. Rashid and Khalid (2017) and Fidora et al. (2007) look at the impact on foreign portfolio investment (FPI). Finally, Nydahl (1999) and Rashid and Waqar (2017) gauge the effect on firms' internationalization and exporting decisions.

A common definition of exchange rate exposure is the sensitivity of firms' market value (stock returns) to unpredictable variations in exchange rates (Dumas, 1978; Adler & Dumas, 1980; Hodder, 1982; Madura, 1989). Sensitivity to the exchange rate is defined as the possible direct loss (mainly due to unhedged exposure) or indirect loss in the firm's cash flows, assets, liabilities and net profit owing to exchange rate variations and, in turn, reductions in its stock market value (Papaioannou, 2006). Foreign exchange exposure is decomposed into the effect of exchange rate uncertainty on the value of net monetary assets with fixed nominal payoffs and the effect of exchange rate volatility on the value of real assets held by the firm.

Unlike domestic monetary assets, both short-term and long-term foreign monetary assets are exposed fully to exchange rate risks. In contrast, the value of real assets may be affected by exchange rate movements regardless of location (Levi, 1994). Thus, besides exchange rate effects on multinational firms, purely domestic firms such as utilities may also be affected by exchange rate movements through effects on aggregate demand or on the cost of traded inputs. Domestic firms that sell goods competing with imports will also be exposed to exchange rate movements (Jorion, 1990; de Jong et al., 2006). While any appreciation in the local currency would adversely affect the value of real assets for multinational firms that rely heavily on exports, it would benefit those multinationals (in terms of their foreign real assets) producing goods that are imported by the home country (Jorion, 1990). Managing exchange rate risk thus becomes a central part of decision making and activity for firms exposed to foreign currency risks (Allayannis et al., 2001).

Hedging strategies to eliminate or at least reduce exchange rate exposure require understanding how, when and to what extent the volatility of the exchange rate affects firm operations, hedging techniques and the instruments available to deal with exchange rate exposure (Barton et al., 2002). For nonfinancial firms, managing exchange rate exposure is independent of their core business and generally handled by corporate treasuries, whereas multinational firms have risk committees to oversee the treasury's strategies for exchange rate exposure (Lam, 2003). This shows that foreign exchange exposure and its management are very important for both domestic and multinational firms.

There are three main types of exchange rate exposure: transaction exposure, translation exposure and economic exchange rate exposure (see Shapiro, 1996; Madura, 1989; Rashid, 2010). Transaction exposure involves cash flow risk and refers to the effect of exchange rate movements on transactional accounts related to receivables (export contracts), payables (import contracts) or repatriation of dividends (Papaioannou, 2006). Exchange rate variations in the currency of denomination of such contracts will result in direct transaction exchange rate exposure for firms.

Translation exposure is balance sheet exchange rate exposure, relating exchange rate movements to the valuation of a foreign subsidiary and, in turn, to its consolidation with the parent company's balance sheet (Shapiro, 1996; Madura, 1989). Translation exposure for a foreign subsidiary is generally measured by the exposure of net assets to expected exchange rate movements. In consolidating financial statements, the translation can be done either at the average exchange rate for that period or at the end-of-period exchange rate, depending on the accounting regulations affecting the parent company. Income statements are generally translated based on the average exchange rate for that period, but the balance sheet exposure of foreign subsidiaries is often translated at the prevailing current exchange rate at the time of consolidation (Papaioannou, 2006).

Economic exchange rate exposure refers to the sensitivity of a firm's present value of future operating cash flows to exchange rate variations. Economic exposure is related to the effect of exchange rate changes on operating expenses (cost of domestic inputs and imports) and revenues (domestic sales and exports). Economic exposure is usually applied to the present value of future cash flows from the operations of a firm's parent company and foreign subsidiaries. Thus, economic exchange rate exposure directly affects the firm's market value. Economic exchange rate exposure is long-lasting, with far-reaching effects on firms'

value. Given this, one can define economic exposure as the effects of unexpected fluctuations in exchange rates on firms' market value. Put differently, a firm would be said to suffer from economic exchange rate exposure if its market value (share price) was significantly influenced by unpredictable variations in exchange rates.

This paper focuses on quantifying the economic exposure of exchange rate changes for a large sample of nonfinancial firms listed on the Pakistan Stock Exchange (PSE), by estimating the effect of exchange rate changes and volatility on their stock returns. Section 2 presents a literature review. Section 3 provides the study's methodology. Section 4 gives the variable definitions and data. Our empirical results are discussed in Section 4. Section 5 presents the study's conclusions.

2. Literature Review

The three approaches most commonly applied when exploring the relationship between exchange rates and stock returns are the traditional approach, portfolio approach and asset approach.

According to the traditional approach, exchange rate changes lead stock prices. For instance, currency depreciations increase the competitiveness of firms' exports, which, in turn, changes the value of their assets and liabilities, culminating in higher profits and hence higher stock prices – at least in the short run (Kutty, 2010). Solnik (1987), who originally proposed this channel, argues that real currency appreciation is bad news for a domestic corporation because it reduces its competitive ability to export. On the other hand, a real depreciation improves domestic firms' ability to export in the short run.

The portfolio approach asserts that exchange rates are determined by a market mechanism and, therefore, stock prices are expected to lead the exchange rate with a negative correlation. This is because a reduction in stock prices reduces domestic wealth, which lowers the demand for money and, in turn, interest rates. A decrease in domestic stock prices also reduces foreign investors' demand for domestic real assets and domestic currency. Such changes in currency demand and supply result in capital outflows and depreciation of the domestic currency (Granger et al., 2000; Caporale et al., 2002; Stavárek, 2005; Pan et al., 2007; Aydemir & Demirhan, 2009; Kutty, 2010).

Finally, under the asset approach, the current money price of any asset is equivalent to its discounted future money prices. This approach

considers the exchange rate to be the price of an asset, that is, the price of one unit of foreign currency in domestic currency. Like the prices of other assets, exchange rates are derived from expected future exchange rates. Any news or factor that affects the future value of the exchange rate will also affect today's exchange rate. Such news or factors may be different from those that influence stock prices. In such a scenario, there would be no relationship between exchange rates and stock prices (Muhammad & Rasheed, 2002). Different approaches predict different relationships between exchange rates and stock prices.

In addition to these three approaches, the literature offers other explanations for firms' exchange rate exposure and their market values. If investors in stock markets are either risk-neutral or risk-seeking, then the exchange rate risk may be viewed as an additional chance to earn higher stock returns. This could increase the demand for stocks that are more exposed to exchange rate risk. On the other hand, risk-averse investors will consider the exchange rate risk an extra cost and thereby invest less in stocks of firms that are more exposed to exchange rate variations. In this context, the effect of exchange rate exposure on the value of firms is conditional on the attitude of investors toward risk. Therefore, we argue that the theoretical literature does not necessarily offer a consensus on the nature of the relationship between exchange rates and stock prices, implying that the effect of exchange rates and exchange rate volatility on stock returns is an empirical issue.

On reviewing the empirical literature, we find that firms' exposure to foreign exchange rates has been studied extensively in the context of developed countries. For example, Choi and Prasad (1995) examine the effect of exchange rates on US firms, He and Ng (1998) study exchange rate effects for Japanese firms, and Al-Shboul and Anwar (2014) provide evidence for Canadian firms. These empirical studies mostly find significant exposure in the case of multinational and domestic firms.

Empirical studies also provide conflicting evidence of the association between exchange rates and firm value (stock prices). For example, Jorion (1990, 1991), Bartov and Bodnar (1994), Choi and Prasad (1995), Allayannis (1996) and Doidge et al. (2000) report no significant association between exchange rate changes and firm value. However, other studies present evidence that the exchange rate has a significant and strong influence on firms' market value and stability (see, for instance, Hodder, 1982; Miller & Reuer, 1998; Bodnar & Wong, 1999; Marston, 2001; Pritamani et al., 2004). Similarly, Baum et al. (2001) show that a decline in the domestic currency value decreases firm profits.

In a recent study, Rashid and Waqar (2017) use a sample of Pakistani nonfinancial firms to show that real exchange rate changes are positively related to firms' export performance, whereas real exchange rate volatility is negatively related to their export performance. Analogously, Rashid and Khalid (2017) investigate the sensitivity of FPI to exchange rate risk and find that exchange rate changes and volatility both have negative, significant and long-lasting effects on FPI in Pakistan.

The theoretical rationale above indicates that domestic companies that are not engaged in international sales may also be affected by exchange rate instability due to competition with other companies (Jorion, 1990; de Jong et al., 2006). Empirically, there is mixed evidence of the impact of exchange rate risk on domestic firms. For example, Jeddi (2001) reports that domestic firms are not exposed to exchange rate changes to any considerable degree. There is also evidence that exchange rate exposure has greater influence on multinational firms' profitability compared to their domestic counterparts (Doukas et al., 2003). On the other hand, Aggarwal and Harper (2010) show that domestic firms' value as well as profitability is significantly exposed to exchange rates and that the effect of exchange rate volatility on domestic firms is not much different from the exchange rate effect on multinational firms.

However, the literature on the effect of exchange rates on firms' value in developing countries is relatively limited (Rashid, 2010). Exchange rate variations are generally more visible in highly inflationary developing economies. For instance, during 1991–98, Turkey experienced high inflation and rapid currency depreciation. For this period, Kiymaz (2003) compares the variance in exchange rate movements of the Turkish lira in terms of other currencies (such as the US dollar, pound sterling, French franc, and German mark) and the variation in foreign currencies in terms of other foreign currencies. His analysis indicates that exchange rates involving the Turkish lira were the most volatile.

Like other developing countries, Pakistan has experienced a series of high-inflation and quick-currency-depreciation episodes over the last few decades. The energy crisis, political instability and disparities in the balance-of-payment accounts have also contributed significantly to rapid fluctuations in the exchange rate. Further, like other developing and emerging economies, firms operating in Pakistan might not actively hedge their exchange rate exposure. Given this background, the economic exchange rate exposure of Pakistani firms is worth exploring, both from a policy point of view as well as from the perspective of investors and firm managers.

Many studies have looked at the impact of exchange rates on different variables in Pakistan, focusing particularly on exchange rate determinants, exchange rate stability and its relationship with different variables such as trade, interest rates, the rate of inflation, exports, FPI and financial derivatives.[1] We depart from such studies in that we examine the economic exchange rate exposure of nonfinancial Pakistani firms.

In principle, we would expect firms to be more exposed to exchange rate risks in a weak currency environment, such as that prevailing in Pakistan. In general, a modest decline in the local currency benefits firms operating in that country, but when the decline is severe or protracted, several challenges arise for such firms (J.P. Morgan, 2017). A severe depreciation is typically a sign of weak economic fundamentals and is generally associated with falling economic growth. Due to expensive imports, a weaker currency may lead to inflationary pressure in the economy. To protect the currency, the government generally raises interest rates and may introduce other regulatory measures and capital controls. Decreases in the interest rate may further slow the process of growth and capital controls can increase risk for both domestic and multinational corporations (MNCs). Domestic firms with liabilities in strong currencies but cash flows in their home currency are typically under more pressure and may reduce their investments.

This paper provides an opportunity to observe whether the effects of exchange rate changes on firm value are consistent with developed economies. We also examine whether domestic firms and MNCs respond differently to exchange rate changes, and if exchange rate appreciations and depreciations have different effects on firms' stock returns. To carry out the empirical analysis, we use an unbalanced panel dataset for a large sample of nonfinancial firms listed on the PSE over the period January 2000 to June 2014. To mitigate the problem of heteroskedasticity, we use a generalized least squares (GLS) estimator.

The paper is motivated by the deficiency of the literature pertaining to Pakistan on the effect of exchange rate changes on firm value, and the importance of this topic in the international financial literature in general and – given its weak currency environment – for Pakistan in particular. Rashid (2010) is an exception in that he examines economic exchange rate exposure in Pakistan using industry-level stock returns and covering all

[1] See Rashid and Khalid (2017), Rashid and Waqar (2017), Rashid (2013), Kemal (2005), Afza and Alam (2011), and Khan and Qayyum (2007).

types of industries (22 in all) instead of focusing on export-oriented or/and import-oriented industries, over the period June 1991 to December 2007. His findings indicate significant exposure to exchange rates for Pakistani industries.

This paper differs from Rashid (2010) in several ways. First, we use firm-level data. Second, we examine the effects of exchange rate volatility over and above the effects of exchange rate changes, whereas Rashid focuses on the effects of exchange rate fluctuations. Third, we examine the effects of exchange rate changes and volatility on the market value of both domestic firms and MNCs. Finally, and more important, we explore the differential effects of exchange rate appreciations and depreciations on firms' stock returns. Our study also differs considerably from Jorion (1990). We analyze the exchange rate exposure of nonfinancial firms listed on the PSE, while Jorion looks at US multinationals in this context. Moreover, unlike Jorion, we assess the difference in exchange rate exposure across domestic and multinational firms, and examine how the differential effects of appreciations and depreciations vary across domestic firms and MNCs.

Thus far, the literature has not analyzed the effects of exchange rate changes on the value of Pakistani nonfinancial firms. There is also a lack of empirical evidence on whether domestic and multinational firms respond differently to exchange rate variations. Examining the exchange rate exposure of domestic and multinational firms in a relatively weak currency environment such as Pakistan would increase our understanding of the effects of exchange rates on firm value. In this context, our study helps extend the literature.

Our findings could help firm managers design an appropriate strategy for hedging exposure to exchange rates. There are several risk management theories that analyze the effect of risk-hedging decisions on firm value. For instance, Smith and Stulz (1985) and Leland (1998) assess the effect of hedging on debt capacity, while Allayannis and Weston (2001) and Carter et al. (2006) investigate its effect on value. Froot et al. (1993) study the effect of hedging on investment policies. They argue that, due to capital market imperfections, the use of derivatives for risk management strategies can affect firm value – for instance, by reducing expected taxes and financial distress costs, mitigating underinvestment and increasing debt capacity to take advantage of debt tax shields.

3. Methodology

To examine the relationship between firm value and exposure to exchange rates, we select the commonly used two-factor model specified by Jorion (1990), modified to accommodate issues related to the relationship between firm value and exchange rates. We begin by estimating Jorion's model and comparing our results with those obtained by other studies that have also estimated this model. The model is defined as follows:

$$R_{it} = \beta_c + \beta_x \, EX_t + \beta_m \, R_{m,t} + \varepsilon_{i,t} \tag{1}$$

where R_{it} is the stock return of firm i at time t, EX_t is the percentage change in the exchange rate variable in month t (the orthogonalized exchange rate change), $R_{m,t}$ is the market return at the end of t, β_c is the constant term, β_x measures firm-level exchange rate sensitivity in excess of the market's exchange rate reaction, β_m is the beta with respect to market returns, and $\varepsilon_{i,t}$ is the error term with a zero mean and constant variance.

In equation (1), the exposure coefficient (β_x) measures the exposure to exchange rate elasticity as the difference between firms' total exposure elasticity and the market's exposure elasticity adjusted by the firms' market beta. Thus, we refer to β_x as the residual exposure elasticity of the firm. Incorporating market returns into the model controls not only for the effect of macroeconomic factors, but also for the market portfolio's own exposure to exchange rates.[2]

Bartov and Bodnar (1994) argue that investors find it difficult to understand the complex relationship between exchange rate variations and firm value. As a result, mispricing may arise in the market, which then affects stock returns. Since investors discover the impact on firm value gradually after new information is released, the exchange rate can affect stock prices with a lag rather than contemporaneously. Bartov and Bodnar refer to this lagged stock price response as the 'delayed-response hypothesis.' They explain this as the inability of financial analysts and equity markets to adequately and precisely forecast the effects of exchange rate changes on firms' earnings during the current period. This, in turn, accounts for the delayed response of stock prices to exchange rate exposure.

[2] Macroeconomic variables – such as the change in the risk-free rate, the market risk premium and investor sentiment that happens to be correlated with the exchange rate – affect the valuation of all firms. To control for macroeconomic influences on realized returns, we include market portfolio returns in the model proposed by Jorion (1990).

The lagged effect suggests that it takes time for the impact of exchange rate changes to be reflected in stock prices, implying market inefficiency. To this 'delayed-response hypothesis' and in line with other studies such as Doukas et al. (2003), de Jong et al. (2006), Rashid (2010), and Rashid and Khalid (2017), we include the one-month and two-month lagged values of the exchange rate:[3]

$$R_{i,t} = \beta_c + \beta_x\, EX_t + \beta_{x1}\, EX_{(t-1)} + \beta_{x2}\, EX_{(t-2)} + \beta_m\, R_{m,t} + \varepsilon_{i,t} \qquad (2)$$

Economic agents are also very interested in exchange rate volatility, which affects not only a given economy, but also the economies of other countries. Policymakers use information on how exchange rate volatility influences different macroeconomic variables. Firm managers consider exchange rate volatility in different models to measure the risk they face. This study also determines how the volatility of the exchange rate (*ver*) affects firm value:

$$R_{i,t} = \beta_c + \beta_x EX_t + \beta_{x1}\, EX_{(t-1)} + \beta_{x2} EX_{(t-2)} + \beta_m R_{m,t} + \beta_{ver}\, ver(\sigma_t^2) + \varepsilon_{i,t} \qquad (3)$$

In principle, both firm managers and investors view positive and negative changes in exchange rates differently. Proponents of behavioral finance suggest that investors behave differently in case of favorable and unfavorable effects of any event. While investors generally want 'downside protection' from adverse effects, they might also be keen to avail the 'upside potential' to maximize their benefits. This makes it worth exploring whether stock prices respond differently to exchange rate appreciation and depreciation.

To examine the differential impact of exchange rate depreciation and appreciation, we classify exchange rate changes by creating two dummy variables. The dummy for exchange rate appreciation (D_t^{app}) takes a value of 1 when the exchange rate change is positive and 0 otherwise. The dummy for exchange rate depreciation (D_t^{dep}) takes a value of 1 when the exchange rate change is negative and 0 otherwise. We then multiply both dummies by the orthogonal exchange rate changes to gauge the differential effect of appreciation and depreciation:

[3] Our choice of a lag length of 2 is based on the literature. However, we also assume that, since investors trade frequently both in the forex and stock exchange markets, the two-month period is enough for stock prices to fully reflect exchange rate information.

$$R_{i,t} = \beta_c + \beta_x^+ EX_t \times D_t^{app} + \beta_{x1}^+ EX_{(t-1)} \times D_{(t-1)}^{app} + \beta_{x2}^+ EX_{(t-2)} \times D_{(t-2)}^{app} + \beta_x^- EX_t \times D_t^{dep} + \beta_{x1}^- EX_{(t-1)} \times D_{(t-1)}^{dep} + \beta_{x2}^- EX_{(t-2)} \times D_{(t-2)}^{dep} + \beta_m R_{m,t} + \beta_{ver}\, ver(\sigma_t^2) + \varepsilon_{i,t} \quad (4)$$

To analyze the effect of exchange rate changes on stock returns, one could use ordinary least squares to estimate the above models for each firm separately. However, the existence of heteroskedasticity – a common feature of monthly and weekly financial time series – impairs the reliability of these estimates. To mitigate the problem of heteroskedasticity, we apply a GLS estimator and run a panel regression.

While numerous studies have debated the appropriate method to measure exchange rate volatility, the literature does not provide any consensus on adopting a generally accepted method. The most commonly used model is the autoregressive conditional heteroskedasticity (ARCH/GARCH) approach. Both the ARCH and GARCH models are used especially to analyze stock prices, inflation rates and exchange rates. A distinguishing feature of these models is that the error variance may be correlated over time, given the phenomenon of volatility clustering. We measure exchange rate volatility by estimating a GARCH model.

4. Data and Variables

Our paper uses monthly data spanning January 2000 to June 2014. We obtain the daily stock price data for nonfinancial firms listed on the PSE and use this to calculate their monthly stock returns. We consider those firms for which at least 10 years' continuous stock price data is available,[4] which yields a sample of 232 nonfinancial firms. The monthly market index data is also taken from the PSE and the monthly exchange rate data in terms of Pakistani rupees per US dollar is obtained from the State Bank of Pakistan's *Statistical Bulletin* (2016 and earlier). Using monthly data to carry out the empirical analysis will yield more robust estimates of economic exchange rate exposure, since both exchange rates and stock prices are high-frequency time series.

[4] We apply this criterion to obtain robust estimates of economic exchange rate exposure. Considering only those firms for which 10 years' stock price data is available allows us to examine the exchange rate exposure of relatively mature firms and for those that have had time to diversify their sales and other economic activities. This restriction does not create the problem of survival bias in our sample since we allow the entry and exit of firms for which more than 10 years' data is available.

To study the exchange rate exposure of both types of firms, we split our sample into domestic and multinational firms, following Jorion (1990), Doukas et al. (2003) and others. Using the foreign sales ratio (foreign sales/total sales) to identify multinational firms, we consider a firm an MNC if its foreign sales ratio is greater than 10 percent. This criterion splits our sample of 232 firms into 117 multinational firms and 105 domestic firms. Ten firms are dropped from the sample due to insufficient information on their foreign sales. The variables used in the regression are stock prices, the orthogonalized exchange rate, the market index and exchange rate volatility (Table 1).

Table 1: Description of variables

Variable	Notation	Description
Stock returns	$R_{i,t}$	Return of firm *i* for period *t*, defined as $R_{i,t} = ln(P_{it}/P_{i(t-1)})$, where P_t and P_{t-1} are the current and one-period lag stock prices of the *i*th firm.
Percentage change in exchange rate	EX_t	Percentage (orthogonalized) change in the exchange rate in month *t*.
Market return	$R_{m,t}$	The log of the monthly KSE 100 index is used to calculate monthly stock market returns.
Exchange rate volatility	$ver(\sigma_t^2)$	Calculated using a GARCH model.

Note: Market and exchange rate risk factors may be correlated with, or influenced jointly by, certain external shocks. To avoid any bias due to factor correlations, we orthogonalize the exchange rate risk factor and regress exchange rate changes on the market factor to obtain orthogonal residuals and use them in our models.
The selection criterion leads to GARCH (1,1) from which the GARCH variances are extracted.

5. Empirical Results

This section presents the study's results. We start with an analysis of all firms followed by analysis of domestic and foreign firms.

5.1. Exchange Rate Exposure for All Firms

The results for exchange rate exposure and stock prices are summarized in Table 2. Using the change in the orthogonalized exchange rate as the independent explanatory variable in Model 1, we find that, overall, nonfinancial firms are negatively affected by exchange rate changes, but to a statistically insignificant degree. This indicates either that the stock prices of Pakistani firms are not sensitive to contemporaneous orthogonalized exchange rate changes or that nonfinancial firms have the same exchange rate exposure as the market portfolio. However, when we

add two-month lagged responses to orthogonalized exchange rate changes in Model 2, the negative coefficient on exchange rate changes increases in magnitude and becomes statistically significant. This negative exposure effect implies that a depreciation (appreciation) in the Pakistani rupee decreases (increases) firm value.

Our results indicate that lagged exchange rate changes are statistically significant when applying a one-month lag, but statistically insignificant in the case of a two-month lag. This lagged response could reflect the difficulty investors face in understanding the complex relationship between exchange rate variations and firm value (Bartov & Bodnar, 1994). These lagged responses indicate that it takes time for the impact of exchange rate changes to be reflected fully in stock prices, suggesting that the equity market is inefficient when it comes to incorporating all the available information instantly (de Jong et al., 2006). A possible explanation for the positive one-period lagged effect of the exchange rate on firm value is that firms may need time to design strategies to minimize exchange rate exposure and to reap the advantages of a depreciation of the home currency (increased exchange rate).

Table 2: Exchange rate exposure for all firms

Coefficient	Model 1	Model 2	Model 3
β_x	-0.0001	-0.0023	-0.0021
	(0.915)	(0.003)	(0.005)
β_{x1}		0.0037	0.0037
		(0.000)	(0.000)
β_{x2}		0.0005	0.0047
		(0.477)	(0.553)
β_m	0.3768	0.3692	0.3642
	(0.000)	(0.000)	(0.000)
β_{ver}			-0.0008
			(0.005)
β_c	0.0012	0.0033	-0.008
	(0.08)	(0.000)	(0.060)
Number of firms	232	232	232

Note: Values given in parentheses are p-values.
Source: Authors' calculations.

The estimates yielded by Model 3 show that the exposure coefficient on orthogonalized changes in the exchange rate is negative and statistically significant, and that the one-month lagged response and exchange rate volatility are also statistically significant. Taken together, our results indicate

that the current level of the exchange rate significantly and negatively affects nonfinancial firms in Pakistan. However, the one-period lagged exchange rate has a positive and significant effect on firms' market value.

The significance of exchange rate exposure for firm value is consistent with Doukas et al. (1999) for Japanese firms, Kiymaz (2003) for Turkish firms and Rashid (2010) for firms in Pakistan. There are several explanations for the significant effect of exchange rate changes on firm value. First, Pakistan is a small open economy and highly integrated, economically and financially, with other emerging and developed economies. Second, firms operating in Pakistan not only export their products, but also use several imported inputs in production. This implies that changes in the exchange rate would affect them significantly through the impact on operating expenses (the cost of domestic inputs and imports) and revenues (domestic sales and exports). Third, firms might not be actively adopting hedging strategies to counter their exposure to exchange rates, at least during the current month, for the effect of exchange rate changes to become apparent.

5.2. Exchange Rate Exposure for Domestic Firms

The monthly stock returns of 105 domestic firms for the period January 2000 to June 2014 are regressed jointly on exchange rate changes, lagged exchange rate changes, exchange rate volatility and the market index for Models 1, 2 and 3. The results are given in Table 3.

Table 3: Exchange rate exposure for domestic firms

Coefficient	Model 1	Model 2	Model 3
β_x	-0.0014	-0.010	-0.008
	(0.435)	(0.000)	(0.001)
β_{x1}		0.0097	0.014
		(0.000)	(0.000)
β_{x2}		-0.0042	-0.0024
		(0.085)	(0.314)
β_m	0.2970	0.3169	0.3197
	(0.000)	(0.000)	(0.000)
β_{ver}			-0.0129
			(0.000)
β_c	-0.0004	-0.0003	-0.1733
	(0.855)	(0.920)	(0.000)
Number of firms	105	105	105

Note: Values given in parentheses are p-values.
Source: Authors' calculations.

The results of the regression for domestic firms indicate that firms experience a statistically insignificant negative impact of exposure to exchange rates in Model 1. In Model 2, which incorporates lagged exchange rate changes, domestic firms have a significant contemporaneous and one-month lagged response, but an insignificant two-month lagged response.

The exposure coefficients in Model 3 reveal a negative and statistically significant impact of contemporaneous exchange rate exposure. The negative relationship between exchange rate changes and firm value implies that a depreciation of the Pakistani rupee against the US dollar leads to a decline in the market value of domestic firms. Further, the estimates indicate that domestic firms experience a significant one-month lagged impact of exposure to the exchange rate. However, this effect appears to be positive. Pakistan's financial markets are not efficient enough to instantly and fully incorporate information on the exchange rate into their stock prices. We also find that exchange rate volatility has a significant, negative effect on the market value of domestic firms. This implies that unexpected exchange rate variations negatively affect domestic firms' operations.

Overall, the results point to a significant association between exchange rate variations and the value of domestic firms. Our results are consistent with those of several other studies – including Aggarwal and Harper (2010) for US firms, Pan and Tang (2012) for Chinese firms, and Alssayah and Krishnamurti (2013) for UAE firms – which find that, as with MNCs, domestic firms are impacted significantly by exposure to exchange rate changes.

The significant exchange rate exposure impacts we present in this study may be a result of the effect of exchange rate variations on net real assets. As mentioned in the introduction, the value of real assets is affected significantly by exchange rate movements and these effects will hold whatever their location (Levi, 1994). Thus, even purely domestic firms may be affected by exchange rate movements through the exchange rate exposure effect on aggregate demand or on the cost of traded inputs. Domestic firms that sell goods competing with imports will also be exposed to exchange rate movements (Jorion, 1990; de Jong et al., 2006).

5.3. Exchange Rate Exposure for MNCs

The monthly stock returns of 117 multinational firms for the period January 2000 to June 2014 are regressed jointly on exchange rate changes, lagged exchange rate changes, exchange rate volatility and the market index

for Models 1, 2 and 3. The results are summarized in Table 4. The results show that multinational firms experience negative and statistically significant impacts of exposure to the exchange rate in all three models. The negative contemporaneous exposure coefficient means that any depreciation of the Pakistani rupee against foreign currency has an adverse effect on the value of multinational firms. The one-period lagged exchange rate change is positively and significantly correlated to firm value, whereas the second-period lagged exchange rate change is negatively and significantly correlated with firm value for MNCs.

As explained above, firms may need time to reap the benefits of a local currency depreciation. However, an ongoing depreciation will adversely affect their activities and, in turn, negatively affect their market value. Another possible explanation for the negative effect of second-period lagged exchange rate changes is that a continuous decrease in the value of currency may adversely affect investors' confidence: if the demand for MNC shares falls, this will have a negative effect on firms' market value. Our findings on the lagged response to exchange rate exposure support Bartov and Bodnar (1994), but are in contrast to the findings of He and Ng (1998) and Nydahl (1999).

Table 4: Exchange rate exposure for MNCs

Coefficient	Model 1	Model 2	Model 3
β_x	-0.0043	-0.0082	-0.0075
	(0.013)	(0.000)	(0.000)
β_{x1}		0.0156	0.017
		(0.000)	(0.000)
β_{x2}		-0.0100	-0.0096
		(0.000)	(0.000)
β_m	0.3124	0.3696	0.370
	(0.000)	(0.000)	(0.000)
β_{ver}			-0.0037
			(0.053)
β_c	0.0004	0.0025	-0.047
	(0.768)	(0.071)	(0.06)
Number of firms	117	117	117

Note: Values given in parentheses are p-values.
Source: Authors' calculations.

In Model 3, the exchange rate volatility variable appears to be significant and negative. This implies that exchange rate volatility has a deteriorating effect on the market value of multinationals – a finding

consistent with that for domestic firms. This suggests that, regardless of whether a firm is domestic or multinational, its value is affected negatively by exchange rate volatility.

Overall, our results for the significant effect of exchange rate variations on the value of multinational firms are consistent with Dominguez and Tesar (2006) and Faff and Marshall (2005). Both studies investigate the effect of exchange rates on multinational firm value for developed and emerging economies (the UK, Canada, Japan and Germany) and obtain significant results. However, our findings contradict Al-Shboul and Anwar (2014) and Bhuiya et al. (2015), who obtain weak results for the exchange rate exposure of Canadian and British multinational firms, respectively. A possible reason for MNCs' significant exposure to exchange rate changes is that they are actively involved in international transactions and might not use foreign currency derivatives extensively to hedge against exchange rate movements.

In sum, the results given in Tables 3 and 4 suggest that the contemporaneous as well as two-month lagged impact of exchange rate changes on the market value of both domestic and multinational firms is significant and negative. Moreover, the two-month lagged response of firms is stronger than the contemporaneous response, while the lagged effect of exchange rate changes is larger for MNCs. In contrast, the one-month lagged response of firm value to exchange rates is, surprisingly, positive and significant.

These findings have several important implications. The significant, lagged firm value response strongly supports the 'delayed response hypothesis' proposed by Bartov and Bodnar (1994). According to this hypothesis, financial experts and investors may be unable to adequately and precisely foresee the effects of exchange rate changes on firms' earnings during the current period. Rather, they discover the impact of exchange rate changes on firms' operations gradually. This inability to effectively and accurately forecast the impact of exchange rates on firms' cash flow streams in the current period could explain the delayed response of stock prices to exchange rate exposure. The findings also suggest that stock prices in Pakistan take time to fully reflect the negative effects of exchange rate exposure.

We offer three alternative explanations for the larger impact of the two-month lagged exchange rate change relative to the contemporaneous exchange rate change. The first is related to forecasting ability: even though

investors and financial specialists might forecast the exchange rate effect on firms' earnings in anticipation of exchange rate changes, they are in a better position to update and modify these forecasts given the realized response to exchange rate changes. The second explanation is that, while financial investors might accurately predict the effect of exchange rate changes before they occur, they still need time to divert their portfolio allocation away from stocks that are highly exposed to exchange rate risk. The third explanation is based on the market efficiency hypothesis. Pakistan's equity market is not efficient enough to fully and instantly incorporate all information on exchange rate movements. This inefficiency results in the delayed response of stock prices to exchange rate changes.

The positive relationship between firm value and the one-period lagged exchange rate is somewhat surprising and adds to the puzzle: why do firms sometimes experience positive impacts of exchange rate changes? A possible explanation is that both domestic firms and MNCs may take time to design strategies to minimize their exchange rate exposure and take advantage of a home currency depreciation (increased exchange rate). Financial hedges, currency netting, finding new foreign markets in which to sell outputs and sourcing factor inputs overseas all require time. It is also possible that firms might sign contracts in advance and in anticipation of exchange rate movements to mitigate adverse exchange rate effects and to take advantage of favorable exchange rate movements. The impact of exchange rate exposure on firm value may vary over time because of these contracts.

Our evidence on the time-varying response of stock prices to exchange rate changes suggests that exchange rate movements discourage investors in the current period. In the following period, they may hedge exchange rate risk using suitable hedging strategies and invest more in exchange rate risk-exposed stocks, perhaps to obtain higher exchange rate risk premiums. However, since they may choose not to bear hedging costs for longer periods and since continuous exchange rate movements give investors a negative signal, they may sell the stocks of firms facing exchange rate exposure. Although we are unable to offer any adequate explanation for the positive impact of the two-month lagged exchange rate, we propose a framework to examine the contemporaneous and lagged firm value response conditional on the direction of exchange rate changes.

5.4. Differential Exchange Rate Exposure

Having established the effects of exchange rate changes on both domestic and multinational firms, we present another set of results to examine the differential impact of exchange rate depreciations and appreciations. We classify exchange rate changes by creating two dummy variables. The dummy for exchange rate appreciations takes a value of 1 when the exchange rate change is positive and 0 otherwise. The dummy for exchange rate depreciations takes a value of 1 when the exchange rate change is negative and 0 otherwise. We then multiply both dummies by the orthogonal exchange rate changes to obtain the differential effect of exchange rate appreciations and depreciations for the whole sample as well as for multinational and domestic firms. The results of the fixed effects model are presented in Table 5. We estimate three models: Model 1 for all firms, Model 2 for domestic firms and Model 3 for multinational firms.

The estimates indicate that exchange rate appreciations and depreciations affect stock returns differently. For example, firms are negatively affected when exposed to current exchange rate appreciations, whereas exchange rate depreciations have a significant positive effect on stock returns. Similarly, a one-period lagged exchange rate appreciation has a positive impact, whereas a one-period lagged exchange rate depreciation has a significant and negative impact on stock returns. Finally, we find that a two-month lagged exchange rate appreciation does not have any statistically significant impact on firm value. However, a two-month lagged exchange rate depreciation is negatively related to stock returns. These findings suggest that current and lagged exchange rate appreciations and depreciations have different effects on firms' stock returns.

The effect of exchange rate appreciations and depreciations also differs across domestic and multinational firms. For instance, although both domestic and multinational firms are negatively affected by current exchange rate appreciations, this effect is statistically insignificant in the MNC case. Similarly, a two-month lagged exchange rate appreciation has a negative effect on MNC stock returns, but we find no such effect for domestic firms. Interestingly, both types of firms are positively and significantly affected by a one-month lagged exchange rate appreciation.

Table 5: Exchange rate depreciation versus appreciation exposure

Coefficient	All firms	Domestic firms	Multinational firms
β_x^+	-0.012	-0.012	-0.002
	(0.000)	(0.000)	(0.387)
β_{x1}^+	0.031	0.016	0.016
	(0.000)	(0.000)	(0.000)
β_{x2}^+	0.002	0.001	-0.007
	(0.475)	(0.987)	(0.002)
β_x^-	0.007	0.004	-0.019
	(0.000)	(0.354)	(0.000)
β_{x1}^-	-0.019	-0.007	0.006
	(0.000)	(0.229)	(0.119)
β_{x2}^-	-0.002	-0.006	-0.018
	(0.074)	(0.113)	(0.000)
β_m	0.361	0.306	0.358
	(0.000)	(0.000)	(0.000)
β_{ver}	-0.235	-0.164	-0.633
	(0.000)	(0.000)	(0.100)
β_c	0.005	0.007	-0.002
	(0.000)	(0.002)	(0.318)
Number of firms	232	105	117

Source: Authors' calculations.

The results given in the table also show that the effects of exchange rate depreciations differ significantly between domestic and multinational firms. Specifically, we find that exchange rate depreciations do not have a significant effect on the stock returns of domestic firms. In contrast, MNC stock returns are affected significantly and negatively by exchange rate depreciations.

6. Conclusion

Exchange rate variations can have a significant effect on firms' market value by affecting their cash flows, assets and liabilities. In this study, we have analyzed 232 firms listed on the PSE for the period January 2000 to June 2014. The results of the all-firms regression suggest negative and significant effects of exchange rate exposure. Specifically, we find that exchange rate changes have a negative and statistically significant effect on the market value of firms. The results also show that the volatility of the exchange rate has a significant and negative effect on firms' market value.

Having examined the exchange rate exposure effects of the whole sample, we separate domestic and multinational firms to establish whether this exposure differs across the two types of firms. Analyzing a sample of 105 domestic firms yields negative and significant exposure coefficients. We also find that the one-month lagged response to exchange rate changes is significant for domestic firms. The effect of exchange rate exposure in this case could be explained as follows: product prices are effectively determined by market forces and the impact of exchange rate fluctuations is reflected in a firm's stock prices and ultimately in its value or profitability.

Next, we analyze the exchange rate exposure of 117 multinational firms, where exchange rate changes have a significant and negative effect on their market value. We also gauge their one-month and two-month lagged responses to exchange rate changes. Our results suggest that the market value of both types of firms is significantly and negatively affected by exchange rate volatility. These findings imply that the exchange rate volatility variable is an important empirical determinant of firm value, regardless of whether the firm is domestic or multinational.

The lagged response of multinational firms to exchange rate changes is stronger than that of domestic firms. One possible reason for this relatively strong effect is that multinationals are more involved in internationalization and exporting activities, which means that their operations and other activities are more exposed to unexpected fluctuations in exchange rates. Our findings suggest that such firms must employ effective and active hedging strategies against changes in exchange rates to maintain their profits and value. Moreover, the government should design policies that help stabilize exchange rates, in turn, enhancing the market value of firms. There is also a need to provide an environment that not only helps, but also encourages, firms to effectively use hedging instruments against exchange rate exposure.

By presenting robust evidence on economic exchange rate exposure for domestic and multinational firms as well as the differential effects of positive and negative exchange rate changes on the market value of both types of firms, we have attempted to contribute to the scarce literature on foreign exchange rate exposure for firms operating in emerging and developing countries such as Pakistan. This analysis could be extended by examining the economic exchange rate exposure of other types of firms, such as exporting versus nonexporting firms, and small versus large firms.

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Market Returns to Education in Pakistan, Corrected for Endogeneity Bias

Sajjad Haider Bhatti*, Muhammad Aslam and Jean Bourdon*****

Abstract

This paper estimates the Mincer wage model for Pakistan's labor market, using a relatively recent dataset and new independent variables. We employ instrumental variables and two-stage least squares to address the problem of the endogeneity of education. Our results show that the returns to education are biased downward due to endogeneity, with significant wage gaps emerging among different regions, between genders and between urban and rural job markets. The study's choice of instruments has conceptual as well as empirical grounds. Our findings establish that the wage determination process is different for males and females across provincial labor markets.

Keywords: Endogeneity of education, human capital model, instrumental variables, Mincer regression, labor market, returns to schooling, Pakistan.

JEL classification: C26.

1. Introduction

The labor market remunerates workers based on their knowledge and the competencies they have acquired over their working life. Collectively, these skills comprise a worker's human capital. The interrelation between human capital and market earnings is captured by the human capital earnings function developed by Mincer (1958, 1974), Becker (1975) and Ben-Porath (1967). This model shows how formal education and on-the-job training (experience) are the key factors contributing to human capital. Mincer's (1974) model of human capital is commonly used to determine the impact of education and training on remuneration (see, for instance, Willis, 1986; Card, 1999). Even in its most basic form, the model

* Department of Statistics, Government College University, Faisalabad, Pakistan.
** Department of Statistics, Bahauddin Zakariya University, Multan, Pakistan.
*** IREDU- CNRS, Université de Bourgogne, Dijon, France.

remains a key approach to calculating earning differentials associated with urban versus rural status, unionized versus non-unionized status, gender, ethnicity, public versus private sector work, and religion (see Johnson & Chow, 1997; Kimmel, 1997; Lassibille, 1998; Korsun, 2010).

Despite the extensive application of the Mincer model, simple estimations using ordinary least squares (OLS) may be biased due to endogenous schooling variables and sample selectivity (Griliches, 1977, 1979; Angrist & Krueger, 1991). The endogeneity problem arises when the explanatory variables are correlated with the error terms: this problem is built into the Mincer earnings function because all individuals in a wage market can never be exactly the same with respect to different unobservable and observable characteristics (Bhatti et al., 2013). Innate ability, for instance, is treated as a determinant of wages, but it is also correlated with schooling. Since innate ability is not directly measureable, simple OLS estimations are likely to yield biased results (Griliches, 1977; Card, 2001), with adverse implications for policymaking.

This endogeneity bias is generally addressed by estimating the human capital earnings function using an approach that combines instrumental variables with two-stage least squares (IV-2SLS), developed by Theil (1953) and Basmann (1957). The technique relies on different instruments of exogenous variation, including birth cohorts, the education level of other household members, IQ test scores, changes in education system and household sibling composition (Ashenfelter et al., 1999; Butcher & Case, 1994).[1] The literature suggests that the schooling coefficient is higher when one accounts for possible endogeneity using IV-2SLS relative to simple OLS.

While many studies have estimated a human capital earnings function for Pakistan, most have not tackled econometric problems such as the potential endogeneity of the schooling variable.[2] The few studies that do so (using instrumental variables) include Abbas and Foreman-Peck (2008), Aslam et al. (2008) and Aslam (2009). This points to ample room for quantifying the economic returns to education in the Pakistani labor market, corrected for endogeneity bias using new instruments. The rest of the article is structured as follows. Section 2 describes the data, variables and method used, including the nature of the instrumental variables involved. Section 3 presents the results of the Mincer model, estimated using OLS and IV-2SLS.

[1] For further reading, see the references given in Bhatti et al. (2013).
[2] See, among others, Qureshi (2012); Aslam and Kingdon (2009); Farooq and Sulaiman (2009); Hyder and Reilly (2005).

The model is estimated separately for males and females and for the four provinces. Section 4 briefly discusses these results relative to the literature on Pakistan and Section 5 concludes the paper.

2. Methodology

We have used data from the Labor Force Survey conducted by the Pakistan Bureau of Statistics in 2008/09. After cleaning the data, we are left with 19,574 observations. We take the monthly wage of individuals as the response variable. The independent variables include different human capital and other social, economic and regional factors contributing to wage determination (Table 1).

Table 1: Variables used to estimate Mincer earnings function

Variable	Definition
Dependent variable	
Lnwage	Natural logarithm of the individual's monthly wage (main job)
Independent variables	
Education	Number of years of schooling completed
Experience	Potential experience measured in years (age – 6 – SCH1)
Experience2	Potential experience squared
Hours	Number of hours accounting for monthly salary (hours worked per month)
Gender	Gender (male = 1, female = 0)
Urban	Dummy variable indicating where the individual's work is located (urban = 1, rural = 0)
Professional	Dummy variable indicating whether the individual has a professional or general degree/diploma (professional = 1, general = 0)
Public	Dummy variable indicating whether the individual works in the public or private sector (public = 1, private = 0)
Tmp.contract	Dummy variable indicating that the individual has temporary employment (without contract) (reference category)
Fix.contract	Dummy variable indicating whether the individual has a fixed-term contract (yes = 1, otherwise 0)
Perm.contract	Dummy variable indicating whether the individual has a permanent contract (yes = 1, otherwise 0)
Punjab	Dummy variable indicating that the individual is from Punjab (reference category)
Sindh	Dummy variable indicating whether the individual is from Sindh (yes = 1, otherwise 0)
KP	Dummy variable indicating whether the individual is from KP (yes = 1, otherwise 0)
Balochistan	Dummy variable indicating whether the individual is from Balochistan (yes = 1, otherwise 0)

Identifying a relevant valid instrument is difficult (Bound et al., 1995). While many studies have used different instruments in IV-2SLS estimations of the earnings function, the issue of endogeneity bias persists (Angrist & Krueger, 1992). We estimate the Mincer earnings function for Pakistani wage market data, using two instrumental variables.

- Instrument 1 is the average numbers of years of schooling attained in the enumeration block where the individual's household is located. The rationale for using this instrument is to combine the effects of several commonly used exogenous factors, including distance to school, general trend toward education, and social environment.
- Instrument 2 follows a similar logic, but also accounts for other factors at the time the individual leaves formal schooling and joins the workforce. Following Bhatti et al. (2013), we define Instrument 2 as the average number of years of schooling attained by a given gender and age group in the year the individual joined the labor market. The data is taken from Barro and Lee (2010), who have estimated the average number of years of schooling for a large sample of countries for 1950–2010.[3]

3. Data Analysis and Results

Tables 2–5 give the results of the estimated human capital earnings function for the Pakistani labor market. The first column in each case presents the OLS results (uncorrected for potential endogeneity bias), while the remaining columns present the IV-2SLS estimation results using both instrumental variables. In the first-stage regressions, the substantially low p-values (given in parentheses) linked to the instrumental variables indicate that both can be considered relevant instruments. The validity of these instruments is assessed using the Sargan test, which indicates that they are valid for Balochistan, but only for females in the other three provinces.

The analysis employs separate Mincer wage models for males and females in all four provinces. Tables 2–5 give the OLS and IV-2SLS estimation results for males and females in Punjab, Sindh, Khyber Pakhtunkhwa (KP) and Balochistan, respectively. In all four cases, the Hausman (1978) exogeneity test statistic indicates significant bias due to the endogeneity of the schooling variable. A comparison of the OLS and IV-2SLS results for endogenous schooling reveals a downward bias associated with this endogeneity. The downward bias in returns to

[3] These measures of average schooling by gender and age group for different countries are available at www.barrolee.com.

schooling is higher for males in Sindh and KP, and for females in Punjab, but similar for both in Balochistan.

The higher coefficient in the IV-2SLS estimations is in line with international studies such as Butcher and Case (1994), Card (1999), Ashenfelter et al. (1999) and Chen and Hamori (2009) as well as Pakistan-based studies such as Abbas and Foreman-Peck (2008) and Aslam (2009). Bhatti et al. (2013) find similar differences in endogeneity-corrected and uncorrected returns to schooling, using similar instruments applied to data for France.

Table 2: OLS and IV-2SLS estimation results, Punjab

	Males			Females		
Variable	**OLS**	**First stage**	**Second stage**	**OLS**	**First stage**	**Second stage**
Intercept	8.0639	-5.2761	7.7622	8.3405	30.0627	7.6287
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
Education	0.0507	–	0.0755	0.0504	–	0.085
	(<0.0001)		(<0.0001)	(<0.0001)		(<0.0001)
Experience	0.0308	0.0235	0.0366	0.0299	-0.7469	0.0381
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
Experience^2	-0.0004	0.0002	-0.0004	-0.0005	0.0054	-0.0004
	(<0.0001)	(<0.0001)	(<0.0001)	(0.0046)	(<0.0001)	(0.0104)
Hours	0.0011	-0.0183	0.0018	-0.0015	-0.1073	0.0039
	(0.2055)	(0.0028)	(0.05023)	(0.6248)	(<0.0001)	(0.2630)
Urban	0.1115	-0.1426	0.1089	-0.0129	-0.0084	-0.0194
	(0.0016)	(0.5558)	(0.0025)	(0.9042)	(0.9898)	(0.8614)
Professional	0.5445	2.8599	0.4038	0.8179	2.7708	0.6765
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(0.0009)	(<0.0001)
Public	-0.0912	-0.6382	-0.0638	-0.1849	-1.0436	-0.1368
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(0.0024)
Fix.contract	0.0454	1.4152	-0.0103	-0.0065	1.3017	-0.0775
	(0.1341)	(<0.0001)	(0.7451)	(0.9375)	(0.0103)	(0.3786)
Perm.contract	0.2822	2.7877	0.1709	0.3006	2.5786	0.1553
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(0.0497)
Instrument 1	–	0.6752	–	–	0.5424	–
		(<0.0001)			(<0.0001)	
Instrument 2	–	1.6564	–	–	-3.2193	–
		(<0.0001)			(<0.0001)	
Hausman exogeneity test	–	72.34		–	15.38	
		(<0.0001)			(<0.0001)	
Sargan–Hansen test	–	24.77		–	0.01	
		(0.0002)			(0.9208)	
R^2	0.4087	0.5836	0.384	0.4417	0.7104	0.4045
F/Wald statistic	290.9	530.6	257.1	39.9	111.1	37

Source: Authors' estimates.

Table 2 presents the OLS and IV-2SLS estimation results for males and females working in Punjab. The returns to schooling are 7.55 percent for males and 8.05 percent for females, implying that every additional year spent in school increases wages by 7.76 percent. The returns to labor market experience are slightly higher in the IV-2SLS estimation: each additional year of labor market experience increases monthly wages by 3.6 percent for males and 3.8 percent for females. The negative coefficient of work experience squared indicates that the returns to experience are concave, that is, the rate of increase tends to fall as labor market experience rises.

The coefficient of work location (urban or rural) is significant only for males. Urban male workers in Punjab earn nearly 10.89 percent more than their rural counterparts. Male workers with a professional degree earn 40 percent more than those with a general qualification. The corresponding wage premium for females is 67 percent. The wage loss associated with working in the public sector is about 6.3 percent for males and about 13 percent for females. Finally, the results show that males with permanent employment earn about 17.09 percent more than those on a fixed-term contract. The corresponding wage premium for females is 15.53 percent. The subsequent tables (Table 3–5) reveal similar results.

Table 3: OLS and IV-2SLS estimation results, Sindh

Variable	Males			Females		
	OLS	First stage	Second stage	OLS	First stage	Second stage
Intercept	7.8893	-4.9993	7.2441	8.0007	31.4907	7.2394
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
Education	0.0636	–	0.1136	0.0624	–	0.0981
	(<0.0001)		(<0.0001)	(<0.0001)		(<0.0001)
Experience	0.0363	0.1119	0.0425	0.0099	-0.7253	0.0159
	(<0.0001)	(<0.0001)	(<0.0001)	(0.3424)	(0.0019)	(0.0148)
Experience^2	-0.0004	-0.0018	-0.0004	-0.0089	0.0058	-0.0086
	(<0.0001)	(<0.0001)	(<0.0001)	(0.9238)	(0.0019)	(0.9182)
Hours	0.003	-0.0363	0.0048	0.012	-0.1457	0.019
	(0.0031)	(<0.0001)	(<0.0001)	(0.0051)	(<0.0001)	(0.0002)
Urban	-0.0039	-0.4386	0.0253	-0.3076	-0.2443	-0.3351
	(0.9283)	(0.1440)	(0.5894)	(0.1774)	(0.8565)	(0.1556)
Professional	0.2709	2.372	0.0299	0.6767	2.6867	0.5365
	(<0.0001)	(0.0001)	(<0.0001)	(<0.0001)	(0.0005)	(0.0003)
Public	-0.1057	-0.6558	-0.0523	-0.1052	-1.0217	-0.0324
	(<0.0001)	(<0.0001)	(<0.0001)	(0.0610)	(0.0016)	(0.6061)
Fix.contract	0.0572	0.9634	-0.0344	-0.14	0.2583	-0.163
	(0.0828)	(0.0001)	(0.3433)	(0.3001)	(0.7461)	(0.2437)
Perm.contract	0.2782	2.333	0.0511	0.2883	1.5566	0.1757
	(<0.0001)	(<0.0001)	(0.07691)	(0.0011)	(0.0026)	(0.0747)
Instrument 1	–	0.6687	–	–	0.5857	–
		(<0.0001)			(<0.0001)	
Instrument 2	–	1.7457	–	–	-3.3918	–
		(<0.0001)			(<0.0001)	
Hausman exogeneity test	–	307.450		–	19.620	
		(<0.0001)			(<0.0001)	
Sargan–Hansen test	–	4.926		–	1.461	
		(0.0265)			(0.2268)	
R^2	0.5085	0.6068	0.4210	0.4809	0.6842	0.4453
F/Wald statistic	387.4	519.9	330.7	25.63	53.73	23.29

Source: Authors' estimates.

Table 4: OLS and IV-2SLS estimation results, KP

Variable	Males			Females		
	OLS	First stage	Second stage	OLS	First stage	Second stage
Intercept	7.8986	-4.603	7.5054	7.5098	33.5428	7.2194
	(<0.0001)	(0.0348)	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)
Education	0.0559	–	0.0804	0.0631	–	0.0756
	(<0.0001)		(<0.0001)	(<0.0001)		(0.0014)
Experience	0.0389	0.1598	0.0421	0.0245	-0.6386	0.0258
	(<0.0001)	(0.0010)	(<0.0001)	(0.0564)	(0.0001)	(0.0486)
Experience^2	-0.0004	-0.0027	-0.0004	-0.0012	0.0028	0.0001
	(<0.0001)	(0.0001)	(<0.0001)	(0.887)	(0.2571)	(0.7846)
Hours	0.0002	-0.0734	0.0025	0.012	-0.1979	0.0151
	(0.976)	(<0.0001)	(0.1261)	(0.0329)	(<0.0001)	(0.0406)
Urban	0.0718	-0.3079	0.0672	-0.0155	2.597	-0.0471
	(0.1797)	(0.4642)	(0.2202)	(0.9175)	(0.0103)	(0.7662)
Professional	0.5952	2.5934	0.481	0.3895	0.8641	0.3569
	(<0.0001)	(<0.0001)	(<0.0001)	(0.3004)	(0.73911)	(0.3496)
Public	-0.0351	-1.1535	0.0064	-0.3604	-1.2231	-0.3423
	(0.1518)	(<0.0001)	(0.8127)	(0.0001)	(0.0411)	(0.0003)
Fix.contract	0.2128	1.0571	0.1703	0.2256	-0.5132	0.226
	(0.0003)	(0.0206)	(0.0049)	(0.2325)	(0.6192)	(0.2339)
Perm.contract	0.1111	2.12	0.0294	0.3589	-0.4048	0.3502
	(0.0018)	(<0.0001)	(0.4825)	(0.0042)	(0.0063)	(0.0057)
Instrument 1	–	0.6147	–	–	0.4174	–
		(<0.0001)			(<0.0001)	
Instrument 2	–	2.0966	–	–	-2.9922	–
		(<0.0001)			(0.0003)	
Hausman exogeneity test	–	16.446		–	18.446	
		(<0.0001)			(<0.0001)	
Sargan–Hansen test	–	6.964		–	0.558	
		(0.0083)			(0.4550)	
R^2	0.4308	0.5666	0.4044	0.4521	0.5967	0.4470
F/Wald statistic	113.3	175.9	88.45	13.57	21.75	10.81

Source: Authors' estimates.

Table 5: OLS and IV-2SLS estimation results, Balochistan

Variable	Males			Females		
	OLS	First stage	Second stage	OLS	First stage	Second stage
Intercept	8.5959	-8.7668	8.3309	8.1682	32.515	7.8487
	(<0.0001)	(<0.0001)	(<0.0001)	(<0.0001)	(0.0001)	(<0.0001)
Education	0.0533	–	0.0702	0.0594	–	0.0752
	(<0.0001)		(<0.0001)	(<0.0001)		(<0.0001)
Experience	0.028	0.2843	0.0292	0.0365	-0.9402	0.0416
	(<0.0001)	(<0.0001)	(<0.0001)	(0.0095)	(<0.0001)	(0.0051)
Experience^2	-0.0003	-0.0044	-0.0002	-0.0005	0.0063	-0.0006
	(0.0019)	(<0.0001)	(0.0107)	(0.1438)	(0.1139)	(0.1315)
Hours	-0.0041	-0.0685	-0.0024	-0.004	-0.1585	0.0002
	(0.0321)	(<0.0001)	(0.2383)	(0.4762)	(0.0064)	(0.9743)
Urban	-0.0766	0.5578	-0.091	-0.1714	0.4092	-0.1728
	(0.1466)	(0.1605)	(0.0897)	(0.6141)	(0.9061)	(0.6186)
Professional	0.2428	2.8487	0.1564	0.1079	3.0128	0.0402
	(0.0022)	(<0.0001)	(0.0693)	(0.5476)	(0.0931)	(0.831)
Public	-0.1226	-1.3242	-0.086	-0.0281	2.0174	-0.0422
	(<0.0001)	(<0.0001)	(0.0026)	(0.8468)	(0.1734)	(<0.0001)
Fix.contract	-0.0879	-0.103	-0.1025	-0.2649	0.4209	-0.2753
	(0.0396)	(0.7515)	(0.0184)	(0.1292)	(0.8129)	(0.123)
Perm.contract	0.1042	1.9	0.0551	0.3771	2.6163	0.3077
	(0.0034)	(<0.0001)	(0.1686)	(0.0063)	(0.0514)	(0.0363)
Instrument 1	–	0.5425	–	–	0.6752	–
		(<0.0001)			(0.0001)	
Instrument 2	–	2.6031	–	–	-4.5426	–
		(<0.0001)			(0.0001)	
Hausman exogeneity test	–	7.800		–	17.533	
		(0.0053)			(<0.0001)	
Sargan–Hansen test	–	2.469		–	0.692	
		(0.1161)			(0.4050)	
R^2	0.4101	0.6053	0.3976	0.6348	0.6897	0.6189
F/Wald statistic	100.4	199.3	80.39	14.87	16.89	12.65

Source: Authors' estimates.

Following Bhatti et al. (2013), we attempt to choose the most appropriate instrumental variable of the two we have proposed through a correlation analysis. Tables A1 to A8 (in the Appendix) provide partial correlation matrices (with p-values in parentheses) of the response variable, suspected endogenous variable (schooling) and both instruments. We observe that, in most cases, instrument 2 has a lower correlation with endogenous schooling and a higher correlation with the response variable,

which negates its value as a good instrument. On the other hand, instrument 1 has a reasonably strong correlation with the endogenous variable and a smaller correlation with the response variable, which implies it is a more relevant and appropriate instrumental variable. Keeping these correlation measures in view and its conceptual superiority, we give preference to instrument 1 as being more appropriate to the IV-2SLS estimation. Another advantage is that it is generated from the data itself.

4. Discussion

Since the Hausman (1978) test statistic indicates that schooling is endogenous and we consider the IV-2SLS estimation technique more suitable than OLS, this section discusses the IV-2SLS results in more detail. These results show that returns to schooling are, in most cases, 50 percent higher than the OLS results would indicate, which is in line with the results obtained by Bhatti et al. (2013), Butcher and Case (1994) and Card (1993). Returns to labor market experience are slightly higher in the IV-2SLS estimation, with each additional year of labor market experience causing monthly wages to rise at a decreasing rate.

As with Barnet-Verzat and Wolff (2008) and Araï et al. (1996), we find significant gender wage differentials in the returns to education. This is also in line with other Pakistan-based studies such as Yasin et al. (2010), Hyder (2007) and Khan and Toor (2003). The coefficient of work location (urban or rural) indicates that urban males earn more than their rural counterparts, except in Balochistan. The reverse holds true for female workers. A significant wage gap associated with location emerges for Punjab alone. This finding is similar in direction but smaller in magnitude to that of Farooq and Sulaiman (2009).

Our results show that, when corrected for the endogeneity of schooling, the wage premium of holding a professional degree falls substantially for both genders in all provinces. The IV-2SLS estimation indicates that people with a professional degree earn more than those without, reflecting the results obtained by Hyder (2007) and Khan and Toor (2003). However, this difference is not statistically significant for females in KP and Balochistan. Generally, there is a significant wage loss associated with working in the public sector compared to the private sector for both male and female workers. This finding contradicts Hyder (2007), who reports higher earnings for public sector workers in Pakistan.

We also find that people with permanent employment earn more than those on a fixed-term or temporary contract, with male workers in KP being the exception. Our findings concerning provincial effects are in line with many other studies on the Pakistani labor market, including Farooq and Sulaiman (2009) and Khan and Toor (2003). A seeming discrepancy is that, generally, both specifications yield an R^2 term of near 40 percent, which looks small but is common in cross-sectional studies (see Gujarati, 2007).

5. Conclusion

This study has used two new instrumental variables to correct for the possibility of endogeneity bias in schooling. The results reveal the significance of this endogeneity bias and the relevance of both instrumental variables. We find that instrument 1 – the average number of years of education in the enumeration block where the individual's household is located – is superior on statistical as well as conceptual grounds.

The results show that each additional year of schooling increases wages by 7.5 percent (for males), 8.5 percent (for females), 11.3 percent (for males), 9.8 percent (for females), 8.4 percent (for males), 7.6 percent (for females), 7.0 percent (for males) and 7.5 percent (for females) in Punjab, Sindh, KP and Balochistan, respectively. Urban wage markets offer some benefits, but not as much in terms of magnitude and significance. There is a significant wage gain for those with a professional degree. Finally, permanent employment is better remunerated than fixed-contract work.

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Appendix

Table A1: Correlation matrix, Punjab (male)

Variable	Lnwage	Education	Instrument 1	Instrument 2
Lnwage	1	0.3979	0.1408	-0.2699
		(<0.0001)	(<0.0001)	(<0.0001)
Education	0.3979	1	0.515	0.4308
	(<0.0001)		(<0.0001)	(<0.0001)
Instrument 1	0.1408	0.515	1	-0.0542
	(<0.0001)	(<0.0001)		(<0.0001)
Instrument 2	-0.2699	0.4308	-0.0542	1
	(<0.0001)	(<0.0001)	(<0.0001)	

Source: Authors' estimates.

Table A2: Correlation matrix, Punjab (female)

Variable	Lnwage	Education	Instrument 1	Instrument 2
Lnwage	1	0.393	0.1504	-0.3321
		(<0.0001)	(0.001189)	(<0.0001)
Education	0.393	1	0.5398	0.449
	(<0.0001)		(<0.0001)	(<0.0001)
Instrument 1	0.1504	0.5398	1	-0.0142
	(0.001189)	(<0.0001)		(0.760481)
Instrument 2	-0.3321	0.449	-0.0142	1
	(<0.0001)	(<0.0001)	(0.760481)	

Source: Authors' estimates.

Table A3: Correlation matrix, Sindh (male)

Variable	Lnwage	Education	Instrument 1	Instrument 2
Lnwage	1	0.4135	0.3036	-0.3212
		(<0.0001)	(<0.0001)	(<0.0001)
Education	0.4135	1	0.4373	0.4758
	(<0.0001)		(<0.0001)	(<0.0001)
Instrument 1	0.3036	0.4373	1	-0.0253
	(<0.0001)	(<0.0001)		(0.141475)
Instrument 2	-0.3212	0.4758	-0.0253	1
	(<0.0001)	(<0.0001)	(0.141475)	

Source: Authors' estimates.

Table A4: Correlation matrix, Sindh (female)

Variable	Lnwage	Education	Instrument 1	Instrument 2
Lnwage	1	0.347	0.2396	-0.2144
		(<0.0001)	(0.0001)	(0.0005)
Education	0.347	1	0.5197	0.3414
	(<0.0001)		(<0.0001)	(<0.0001)
Instrument 1	0.2396	0.5197	1	-0.1253
	(0.0001)	(<0.0001)		(0.0448)
Instrument 2	-0.2144	0.3414	-0.1253	1
	(0.0005)	(<0.0001)	(0.0448)	

Source: Authors' estimates.

Table A5: Correlation matrix, KP (male)

Variable	Lnwage	Education	Instrument 1	Instrument 2
Lnwage	1	0.4739	0.104	-0.3767
		(<0.0001)	(0.0001)	(<0.0001)
Education	0.4739	1	0.4393	0.5378
	(<0.0001)		(<0.0001)	(<0.0001)
Instrument 1	0.104	0.4393	1	-0.1136
	(<0.0001)	(<0.0001)		(<0.0001)
Instrument 2	-0.3767	0.5378	-0.1136	1
	(0.0001)	(<0.0001)	(<0.0001)	

Source: Authors' estimates.

Table A6: Correlation matrix, KP (female)

Variable	Lnwage	Education	Instrument 1	Instrument 2
Lnwage	1	0.3524	0.1289	-0.4564
		(<0.0001)	(0.1089)	(<0.0001)
Education	0.3524	1	0.3699	0.5059
	(<0.0001)		(<0.0001)	(<0.0001)
Instrument 1	0.1289	0.3699	1	-0.1005
	(0.1089)	(<0.0001)		(0.2125)
Instrument 2	-0.4564	0.5059	-0.1005	1
	(<0.0001)	(<0.0001)	(0.2125)	

Source: Authors' estimates.

Table A7: Correlation matrix, Balochistan (male)

Variable	Lnwage	Education	Instrument 1	Instrument 2
Lnwage	1	0.4913	0.0723	-0.3426
		(<0.0001)	(0.0088)	(<0.0001)
Education	0.4913	1	0.4365	0.5632
	(<0.0001)		(<0.0001)	(<0.0001)
Instrument 1	0.0723	0.4365	1	-0.0856
	(0.0088)	(<0.0001)		(0.0019)
Instrument 2	-0.3426	0.5632	-0.0856	1
	(<0.0001)	(<0.0001)	(0.0019)	

Source: Authors' estimates.

Table A8: Correlation matrix, Balochistan (female)

Variable	Lnwage	Education	Instrument 1	Instrument 2
Lnwage	1	0.5315	0.2008	-0.394
		(<0.0001)	(0.0654)	(0.0002)
Education	0.5315	1	0.4127	0.5087
	(<0.0001)		(<0.0001)	(<0.0001)
Instrument 1	0.2008	0.4127	1	-0.1347
	(0.0654)	(<0.0001)		(0.2192)
Instrument 2	-0.394	0.5087	-0.1347	1
	(0.0002)	(<0.0001)	(0.2192)	

Source: Authors' estimates.

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Financial Development and Output Volatility: A Cross-Sectional Panel Data Analysis

M. Tariq Majeed* **and Ayesha Noreen****

Abstract

This paper aims to provide a more comprehensive understanding of the impact of financial developments on output volatility. Using cross-sectional and panel datasets for 79 countries from 1961 to 2012, we find that financial expansion plays a significant role in mitigating output volatility, although the evidence is weak in some cases. The role of financial stability is more prominent than that of other measures of financial growth in mitigating output volatility. The volatility of terms of trade and inflation contributes positively to increasing output volatility. We also evaluate the channels through which financial developments can affect output volatility. Our model investigates the link between financial growth and output volatility through two potential channels, using four measures of financial development. The volatility of inflation and of terms of trade are used as proxies for monetary sector and real sector volatility, respectively. Financial development plays a mixed role in amplifying or mitigating output volatility through real and monetary sector volatility. Overall, there is some evidence to suggest that financial development amplifies monetary sector volatility, but weaker evidence that real sector volatility is reduced by financial development.

Keywords: Output volatility, financial development, panel data.

JEL classification: E30, E51, G20, O16.

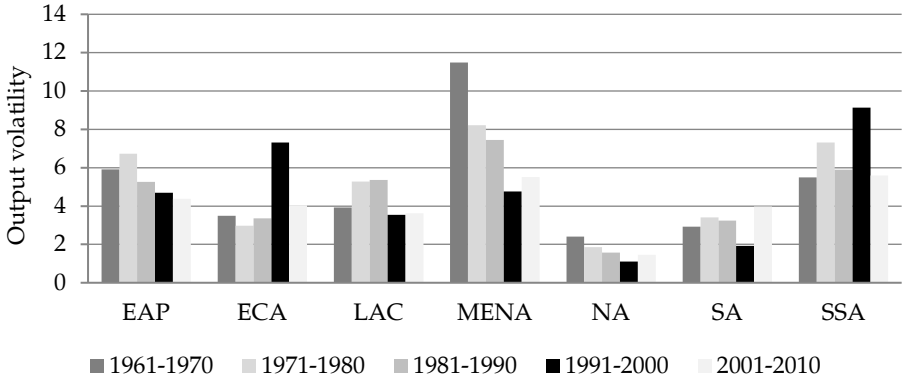
1. Introduction

Over the past few decades, the concept of output volatility has gained greater attention among economists and policymakers. Figure 1 shows that output volatility does not follow a similar pattern across regions. It has declined over time in East Asia and the Pacific, Latin America and the Caribbean, North America, and the Middle East and North Africa. The reverse is true for Europe and Central Asia, sub-Saharan Africa and South Africa, where no clear pattern of output volatility is observable. Figure 2

* Assistant Professor, Quaid-i-Azam University, Islamabad, Pakistan.
** Lecturer, University of Wah, Wah, Pakistan.

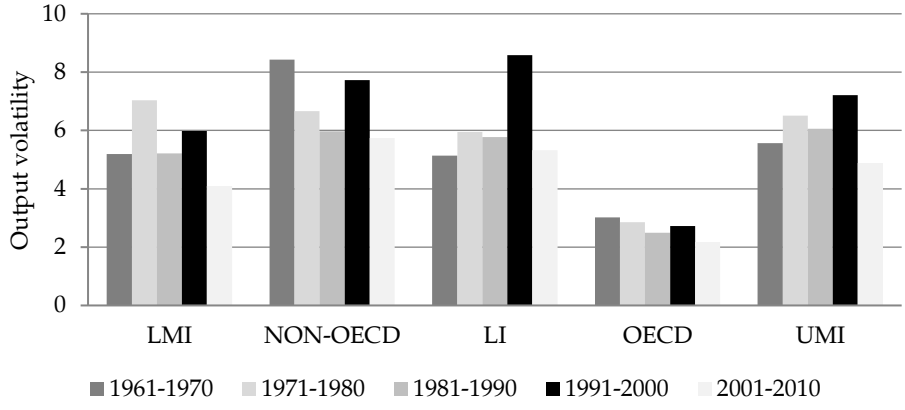
shows that countries with higher levels of economic development are less prone to output volatility. Numerous studies show that high output volatility tends to hinder growth and development (see, among others, Bruno & Easterly, 1998; Loayza & Hnatkovska, 2004; Aghion et al., 2004).

Figure 1: Output volatility across regions, 1961–2010



Note: Output volatility = four-year standard deviation of GDP per capita.
EAP = East Asia and the Pacific, LAC = Latin America and the Caribbean, NA = North America, MENA = Middle East and North Africa, ECA = Europe and Central Asia, SSA = sub-Saharan Africa, SA = South Africa.

Figure 2: Output volatility, by income group, 1961–2010



Note: Output volatility = four-year standard deviation of GDP per capita.
LI = low-income, LMI = lower middle-income, UMI = upper middle-income.

There is broad agreement in the literature that high macroeconomic volatility tends to depress investment, is biased toward short-term returns

and reduces economic growth (Servén, 2002). Recent work shows that higher macroeconomic volatility is also related to lower investment in human capital (Krebs et al., 2005). High output volatility and financial crises are recurring characteristics of many world economies, but they can be serious obstacles to development because they are closely related to high consumption volatility, high poverty, low long-term growth and high inequality. The literature recognizes that financial sector development has a positive impact on economic growth: a well-functioning financial market enables better relationships between investors and savers, promotes diversification, reduces risk, mitigates information asymmetries, encourages individuals to behave more effectively, and helps stabilize the economy and reduce output volatility (Ramey & Ramey, 1995; Aghion et al., 2000).

In recent years, another important area of research has looked at what causes output volatility and how it can be mitigated. The literature examines different determinants of output volatility, including fiscal policy, consumption volatility, remittances, oil prices and foreign direct investment. While financial development is also a key determinant of output volatility, the literature in this area is not substantial and tends to yield ambiguous results. It is difficult to determine whether financial intermediary development leads to an increase or decrease in output volatility or if it helps reduce the impact of external shocks on the economy.

This paper attempts to identify the possible links between financial intermediary development and output volatility by examining whether financial intermediaries help absorb shocks – thereby reducing the effect of real and monetary volatility – or if they intensify the impact of such shocks. To the best of our knowledge, the relationship between output volatility and financial development has not been extensively researched. Therefore, this paper contributes to the literature by exploring the relationship between financial development and output volatility to determine whether financial development magnifies or reduces the volatility of output.

Additionally, the indicators of financial development used in previous studies do not provide a complete picture of the sector. While most studies have used the ratio of private credit to GDP as a proxy for financial development, this indicator does not include other characteristics such as financial depth, stability, access and efficiency. We use four measures of financial development: (i) the ratio of private credit to GDP (financial depth), (ii) net interest margin (financial efficiency), (iii) z-score (financial stability) and (iv) the number of bank branches per 100,000 adults (financial access). Besides looking at the relationship between financial development and

output volatility, we also consider two potential channels through which the sector's development can affect output volatility – real sector volatility and financial sector volatility. Several pre-estimation tests and other diagnostics are used to ensure more reliable results.

The study tests the following hypotheses: (i) that the impact of financial development on output volatility is negative, and (ii) that the impact of financial development on output volatility varies depending on the measure used to gauge financial development. The finance-output volatility nexus may be explained better through the mechanism of shocks to the real and monetary sectors. Therefore, we also test a third hypothesis: that financial development has the power to mitigate or aggravate output volatility through real and monetary sector shocks.

The remaining study is organized as follows. Section 2 presents a literature review of studies on output volatility and financial development. Section 3 illustrates the analytical framework used. Section 4 presents the study's data and variables. Section 5 interprets and discusses the results obtained. Section 6 concludes the study, offering some policy implications.

2. Literature Review

The literature on the finance-and-growth nexus suggests that financial development has diverse effects on economic growth. An increasing number of studies indicate that a well-functioning banking system stimulates economic growth (see Goldsmith, 1969; Levine & Zervos, 1998; Levine et al., 2000; Beck et al., 2000; Beck & Levine, 2004). Other studies disagree: Lucas (1988), for instance, contends that finance is not an important determinant of growth. Similarly, Stern (1989) asserts that finance does not play a role in economic growth. The literature also points to the negative impact of financial development, as suggested by Keynes (1936) and supported empirically by De Gregorio and Guidotti (1995) and Andersen and Tarp (2003). Ram (1999) investigates the relationship between the growth of real GDP per capita and financial development, using panel data for 95 countries over the period 1960–89. The results indicate that financial development has a weak and negative or even negligible impact on the growth of real GDP per capita. Thus, the literature on finance and growth offers mixed results at best.

Although a large body of literature focuses on the growth effects of finance, few studies have examined in detail the potential links between financial development and the volatility of output. Mishkin (2009)

emphasizes that, among other macroeconomic policy objectives, the stability of output growth is also very important. However, theory offers no clear predictions as to the effect of financial development on growth volatility. Bernanke and Gertler (1989), for example, argue that financial constraints to firms can play a vital role in the spread of business cycles, eventually leading to greater variations. On the other hand, studies such as Bacchetta and Caminal (2000), Aghion et al. (2004), Beck et al. (2006) and Ko (2008) argue that the negative or positive effect of financial development on volatility depends on monetary and real shocks, the stages of a country's financial development, and demand shocks or credit supply.

The empirical literature on finance and output volatility also presents mixed findings. For instance, Denizer et al. (2002) find that improved financial systems minimize fluctuations in per capita output growth. Similarly, Bekaert et al. (2006) observe that financial liberalization often results in lower (consumption) growth volatility. Ang (2011) shows that the implementation of financial repressionist policies is strongly associated with lower consumption volatility. Dynan et al. (2006) suggest that financial innovations contributed to stabilization in the mid-1980s and helped reduce output volatility.

Some empirical studies suggest that output volatility increases the effect of financial development. Bernanke and Gertler (1989) find that periods of financial distress correspond to periods of relatively high agency costs in investment. Financial constraints to firms can thus play a key role in the spread of the business cycle and eventually lead to greater variations. Similarly, Aghion et al. (2004) observe that countries undergoing financial development may become more unstable in the short run. Levchenko et al. (2009) find significant evidence to support the view that financial liberalization increases output volatility. Other empirical studies do not find a strong relationship between finance and output volatility. Acemoglu et al. (2003) establish that the volatility effect of financial development diminishes once institutional variables are controlled for. Similarly, Beck et al. (2006) find no robust relationship between financial development and aggregate economic volatility. Thus, the available empirical literature on finance and volatility does not provide a clear picture of this nexus.

The theoretical foundations of this study draw on four strands of the literature. The first, which relates to finance and development, argues that developed financial systems enhance the economy's ability to absorb shocks and help reduce output volatility. Most studies within this area are theoretical. For instance, Aghion et al. (1999) build a macroeconomic model

that combines financial market imperfections with unequal investment opportunities. Their model predicts high output volatility in the absence of a developed financial sector. They argue that savers and investors are separated when the financial sector remains underdeveloped, and credit supply and demand is more cyclical. Investors are likely to be locked out of the credit market when the economy faces an adverse shock and to rush into the credit market when the economy absorbs a positive shock. In such situations, volatility is bound to increase.

Acemoglu and Zilibotti (1997) point out another important link between financial sector development and volatility by highlighting the importance of diversification in reducing risk. They argue that diversification is not possible in the early stages of development, given the indivisibility of capital. However, once wealth accumulates, diversification becomes possible and investment increases: in this situation, investment volatility and risk is reduced. Aghion et al. (2000) assess the role of open economies in explaining the volatility-finance nexus, arguing that volatility increases in open economies with intermediate levels of financial development.

The second strand of the literature explains the link between financial development and volatility through information asymmetries. Notable studies in this area include Bernanke and Gertler (1989), Greenwald and Stiglitz (1993), and Kiyotaki and Moore (1997). These studies develop general equilibrium models to show how asymmetries of information in financial markets exacerbate volatility.

The third strand of the literature highlights the importance of financial development in long-run macroeconomic performance. Studies such as Gertler (1988) and Levine (1997) show that financial development helps reduce the cost of acquiring information and facilitates lower transaction costs. The argument here is that financial development helps reduce information asymmetries, improves corporate governance and leads to better resource allocation. It also enables financial institutions to improve their risk management and allocate their liquidity and funds to the most productive uses (Greenwood & Smith, 1997).

Finally, a fourth strand of the literature cautions against the adverse volatility-related consequences of financial sector development. Shleifer and Vishny (2010) and Wagner (2010) argue that financial development can cause over-leverage or induce greater risk-taking by entrepreneurs and banks, potentially driving up the level of volatility. Many recent studies

show that, while financial intermediaries and institutions can help mitigate friction, amplification mechanisms within the financial sector and from the financial sector to the real sector can exacerbate volatility (for surveys of recent models, see Brunnermeier et al., 2012; Quadrini, 2011).

In sum, the literature demonstrates several potential routes through which financial development can affect output volatility. First, a more developed financial sector is likely to match savers and investors more efficiently, allowing the economy to absorb shocks more easily, thereby reducing volatility. Second, more financial markets and institutions facilitate diversification, which helps reduce risk and volatility. Third, financial development may be a proxy for information asymmetries, which themselves can cause increased volatility. This study contributes to the literature using four diverse indicators of financial performance: depth, efficiency, stability and access. It also presents two mechanisms – real sector and financial sector volatility – through which financial development can amplify or reduce output volatility.

3. Methodology

The global financial crisis has triggered policy debates on the role of finance in proliferating and decreasing macroeconomic fluctuations. The last two decades show that high output volatility can have an adverse effect on economic growth, welfare and poverty, particularly in developing countries (Ramey & Ramey, 1995). The literature discussed in Section 2 outlines various ways in which financial development can affect output volatility. This is the starting point of our analysis.

While there is considerable empirical evidence on the determinants of output volatility, most of these investigations have focused on specific determinants of output volatility, with little uniformity in the sets of explanatory variables used. In recent years, the literature has also looked at how financial openness, financial integration, financial institutions and monetary policy can affect output volatility. Informal arguments suggest that the quality of financial institutions may be a key determinant of output volatility: sound financial institutions enable economies to better handle large output fluctuations.

In this study, we focus on financial development as a key determinant of output volatility. There are contrasting views on the role of financial development in economic growth and output volatility. On the one hand, the development of financial systems changes the incentives and

constraints faced by economic agents by producing information, allocating capital and monitoring firms. Furthermore, financial developments help reduce risk, pool savings and ease exchange, with positive impacts on economic growth (Greenwood & Jovanovic, 1990; Bencivenga & Smith, 1991; King & Levine, 1993; Acemoglu & Zilibotti, 1997). On the other hand, Ko (2008) finds that economies characterized by financial liberalization have higher output volatility in the face of monetary policy shocks. Dynan et al. (2006) suggest that financial innovations contribute to long-term, not short-term, declines in output volatility.

3.1. Output Volatility Model

Following the literature, our benchmark measure of volatility is the standard deviation (SD) of real per capita GDP (see Beck et al., 2000; Hakura, 2007; Malik & Temple, 2009; Ahamada & Coulibaly, 2011). It is also standard in the empirical literature to use the log of GDP per capita as a measure of growth to determine volatility (see Posch, 2011; Beck et al., 2006). Accordingly, we develop our model by introducing the following baseline regression:

$$OV = \alpha_1(lGDP)_{i,t} + \beta_1(FD)_{i,t} + \delta(X)_{i,t} + \mu_i + \varepsilon_{i,t} \quad (1)$$

where OV is macroeconomic volatility, measured by the SD of the annual percentage growth rate of real GDP per capita based on constant 2005 US dollars, and $lGDP$ is a measure of growth, denoted by the log of GDP per capita. FD represents financial development. X is a control variable, trade openness, measuring the ratio of exports plus imports to GDP, following Kose et al. (2009). Finally, μ is a country-specific effect, ε is the error term and i and t denote country and period, respectively.

To evaluate empirical predictions concerning the relationship between financial development and macroeconomic volatility, we would need to construct measures of the ability of financial systems to improve information asymmetries, ease risk management and smooth resource mobilization. Most other studies have used financial depth as the sole measure of financial development (see King & Levine, 1993; Beck et al., 2000; Bacchetta & Caminal, 2000; Denizer et al., 2002). Financial development is, however, a multidimensional concept and warrants more than one measure. A complete picture of financial development and its characteristics is presented by Čihák et al. (2012), who emphasize the importance of these characteristics in cross-county analyses. Our measure of financial development is based on these characteristics and comprises four

components: depth (the ratio of private credit to GDP), efficiency (net interest margin), access (number of bank branches per 100,000 adults) and stability (z-score).

Closely following Beck et al. (2000), we study the responsiveness of macroeconomic volatility to real and monetary shocks. We explore two channels through which financial development and output volatility may be linked. These channels are discussed broadly in the literature as follows. Gavin and Hausmann (1996) show that external terms-of-trade (TOT) shocks are associated with increased growth volatility, while Singer and Edström (1993) and Lutz (1994) find that greater TOT volatility reduces growth. Bleaney and Fielding (2002) show that there is a tradeoff in the choice of exchange rate regime between inflation volatility and stability of output.

Conrad et al. (2010) demonstrate that the relationship between output volatility and inflation volatility is positive, and that the central bank's commitment to reducing inflation volatility helps decrease output volatility. However, the relationship between output and inflation is not that simple: it varies, depending on the type of shock to the economy. In the model developed by Cecchetti et al. (2002), the ratio between inflation and output volatility depends on the slope of the supply curve. Thus, a country that has a relatively high aversion toward inflation variability would suffer more in terms of output volatility.

Introducing these two channels, we estimate the following model:

$$OV = \alpha_1(lGDP)_{i,t} + \alpha_2 SD(TOT)_{i,t} + \alpha_3 SD(INF)_{i,t} + \beta_1(FD)_{i,t} + \delta(X)_{i,t} + \mu_i + \varepsilon_{i,t} \quad (2)$$

The following four regressions are estimated for four distinct measures of financial development:

$$OV = \alpha_1(lGDP)_{i,t} + \alpha_2 SD(TOT)_{i,t} + \alpha_3 SD(INF)_{i,t} + \beta_1(PC)_{i,t} + \delta(X)_{i,t} + \mu_i + \varepsilon_{i,t} \quad (2.1)$$

$$OV = \alpha_1(lGDP)_{i,t} + \alpha_2 SD(TOT)_{i,t} + \alpha_3 SD(INF)_{i,t} + \beta_1(NIM)_{i,t} + \delta(X)_{i,t} + \mu_i + \varepsilon_{i,t} \quad (2.2)$$

$$OV = \alpha_1(lGDP)_{i,t} + \alpha_2 SD(TOT)_{i,t} + \alpha_3 SD(INF)_{i,t} + \beta_1(ZS)_{i,t} + \delta(X)_{i,t} + \mu_i + \varepsilon_{i,t} \quad (2.3)$$

$$OV = \alpha_1(lGDP)_{i,t} + \alpha_2 SD(TOT)_{i,t} + \alpha_3 SD(INF)_{i,t} + \beta_1(BB)_{i,t} + \delta(X)_{i,t} + \mu_i + \varepsilon_{i,t} \quad (2.4)$$

where *SD(TOT)* is the four-year SD of TOT as a measure of real sector shocks and *SD(INF)* is the four-year SD of inflation as a measure of monetary shocks. *PC* is the ratio of private credit to GDP (the financial resources provided to the private sector by domestic money banks) and is used to measure financial depth. *NIM* is the net interest margin (the accounting value of a bank's net interest revenue as a share of its average interest-bearing or total earning assets), measuring financial efficiency. ZS is the bank's z-score (the probability of default of a country's banking system) and measures financial stability. *BB* is the number of bank branches per 100,000 adults and is used to measure financial access.

Despite substantial evidence concerning the role of the financial system in shaping economic development, there are serious shortcomings associated with measuring how this system functions. The literature does not provide sound cross-country, cross-period measures of the degree to which financial systems enhance the quality of information about firms and hence the efficiency of resource allocation, or impose good corporate governance on the firms to which they funnel those resources. Nor are there any measures to indicate whether financial systems provide effective mechanisms to manage, pool and diversify risk; to mobilize savings from disparate savers so that these resources can be allocated to the most promising projects in the economy; or to facilitate trade. Instead, researchers have largely (though not exclusively) relied on measures of the size of the banking industry as a proxy. However, size is not a measure of quality, efficiency or stability. We have therefore used four measures of financial development to obtain a comprehensive picture of the financial sector's role in determining output volatility.

Finally, we also incorporate interaction terms involving financial development and the SD of TOT and that of inflation, as applied by Dabla-Norris and Srivisal (2013) and Beck et al. (2000). We estimate the following regression equation with interaction terms:

$$OV = \alpha_1(lGDP)_{i,t} + \alpha_2 SD(TOT)_{i,t} + \alpha_3 SD(INF)_{i,t} + \beta_1(FD)_{i,t} + \gamma_1(FD * SDTOT)_{i,t} + \gamma_2(FD * SDINF)_{i,t} + \delta(X)_{i,t} + \mu_i + \varepsilon_{i,t} \quad (3)$$

The following four models are estimated for four distinct measures of financial development with interaction terms:

$$OV = \alpha_{11}(lGDP)_{i,t} + \alpha_{12}(vTOT)_{i,t} + \alpha_{13}(vINF)_{i,t} + \beta_{11}(PC)_{i,t} + \gamma_1(PC * vTOT)_{i,t} + \gamma_2 PCvINF_{i,t} + \delta lOPEN_{i,t} + \mu_i + \varepsilon_{i,t} \quad (3.1)$$

$$OV = \alpha_{21}(lGDP)_{i,t} + \alpha_{22}(vTOT)_{i,t} + \alpha_{23}(vINF)_{i,t} + \beta_{21}(NIM)_{i,t} + \gamma_1(NIMvTOT)_{i,t} + \gamma_2 NIMvINF_{i,t} + \delta lOPEN_{i,t} + \mu_i + \varepsilon_{i,t} \quad (3.2)$$

$$OV = \alpha_{31}(lGDP)_{i,t} + \alpha_{32}(vTOT)_{i,t} + \alpha_{33}(vINF)_{i,t} + \beta_{31}(ZS)_{i,t} + \gamma_1(ZSvTOT)_{i,t} + \gamma_2 ZSvINF_{i,t} + \delta lOPEN_{i,t} + \mu_i + \varepsilon_{i,t} \quad (3.3)$$

$$OV = \alpha_{41}(lGDP)_{i,t} + \alpha_{42}(vTOT)_{i,t} + \alpha_{43}(vINF)_{i,t} + \beta_{41}(BB)_{i,t} + \gamma_1(BBvTOT)_{i,t} + \gamma_2 BBvINF_{i,t} + \delta lOPEN_{i,t} + \mu_i + \varepsilon_{i,t} \quad (3.4)$$

where *SD(TOT)* is labelled *vTOT*, *SD(INF)* is labelled *vINF*, and γ_1 and γ_2 are the coefficients of the interaction terms for financial development with the SD of TOT and the SD of inflation, respectively. A negative (positive) sign on γ_1 would indicate that the development of financial intermediaries reduces (magnifies) the incidence of real volatility. A negative (positive) sign on γ_2 would imply that financial intermediaries dampen (aggravate) the incidence of monetary volatility.

3.2. Econometric Techniques

Our analysis spans 79 countries over the period 1961–2012, based on data from the World Development Indicators and International Financial Statistics.[1] We employ dynamic panel data, as the data for GDP per capita depends on its previous values. This enables us not only to control for unobserved cross-country heterogeneity, but also to investigate dynamic relationships.

To examine the relationship between financial development and economic growth, we use the generalized method of moments (GMM) estimators developed for dynamic panel models by Arellano and Bond (1991). Greene (1997) points out that GMM yields consistent and efficient estimates in the presence of arbitrary heteroskedasticity, while Arellano and Bond show that the error term ε is not serially correlated and that the independent variables X are weakly exogenous. With these conditions in mind, they propose a two-step GMM estimator. First, the error terms are assumed to be independent and homoskedastic across countries and over time. Second, the residuals retained in the first step are used to construct a

[1] See http://data.worldbank.org/data-catalog/world-development-indicators for the World Development Indicators and http://www.imf.org/en/Data for the International Financial Statistics.

consistent estimate of the variance-covariance matrix, thereby relaxing the assumptions of independence and homoskedasticity.

3.3. Descriptive Analysis

As Table 1 shows, the minimum value of output volatility is 1.76 for Australia and the maximum value is 10.09 for Rwanda. The minimum value of financial depth (private credit to GDP) is 5.47 for Rwanda and the maximum value is 136.58 for the UK. The minimum value of financial stability (z-score) is –1.40 and the maximum value is 43.08, for Ecuador and Panama, respectively. The minimum value of financial efficiency (the net interest margin) is 0.706 and the maximum value is 12.19, for Ireland and Ghana, respectively. The minimum value of financial access (number of bank branches) is 1.19 and the maximum value is 112.80, for Argentina and Cyprus, respectively. The minimum value of inflation volatility is 1.13 and the maximum value is 58.24, for Germany and Uruguay, respectively. Finally, the minimum value of TOT volatility is 2.26 and the maximum value is 118.48, for Belgium and Iran, respectively.

Table 1: Descriptive statistics

Variable	Mean	SD	Min	Max
Output volatility	4.053454	1.827282	1.762503	10.09317
Growth	8.139591	1.655917	5.026165	10.72645
SD (INF)	4.257263	5.933262	1.129389	58.24967
SD (TOT)	17.15693	18.2415	2.259562	118.4770
Private credit to GDP	41.31117	31.72726	5.466166	136.5836
Bank deposits	49.35951	35.35775	6.907424	163.3433
Net interest margin	4.094926	2.588214	0.706318	12.18597
Overhead cost	3.487886	2.409974	0.3492328	12.86581
Z-score	16.26136	10.54338	-1.40414	43.07984
Bank branches	21.63507	22.91749	1.195648	112.8031
Openness	4.030614	0.5496448	2.827155	5.791479

Source: Authors' calculations

Table 2 shows the correlation among the variables used in the study. Output volatility is positively correlated with volatility in inflation and TOT changes. It is negatively correlated with three financial development indicators – the ratio of private credit to GDP, z-scores and the number of bank branches per 100,000 adults – and positively correlated with net interest margin (financial efficiency). GDP per capita growth is negatively correlated with output volatility, while openness is positively correlated with output volatility.

Table 2: Correlation matrix

Variable	OV	LGDP	SD (INF)	SD (TOT)	PC	NIM	ZS	BB	LOPEN
Output volatility	1.0000								
Growth	-0.4063	1.0000							
SD (INF)	0.0134	0.0927	1.0000						
SD (TOT)	0.2204	0.2228	0.0689	1.0000					
Private credit to GDP	-0.3320	0.7154	0.1278	0.1567	1.0000				
Net interest margin	0.3324	0.6612	0.0999	0.2390	0.6922	1.0000			
Z-score	-0.0713	0.0246	0.0797	0.2610	0.0422	0.0627	1.0000		
Bank branches	-0.2498	0.6436	0.0584	0.1120	0.6306	0.4661	0.0645	1.0000	
Openness	0.0253	0.3573	0.0748	0.1148	0.2147	0.1685	0.1823	0.2252	1.0000

Source: Authors' calculations

4. Estimations and Results

This section presents the study's estimation results.

4.1. Cross-Country OLS Regression Results

Table 3 gives the estimation results for the ordinary least squares (OLS) regression equations given in Section 3, as equations 2.1 to 2.4. These results show that private credit to GDP (column 1) and z-score (column 3) have negative signs, indicating that an increase in financial depth and financial stability leads to a decrease in output fluctuations. This implies that well-developed financial systems can dampen output volatility by removing or alleviating financial constraints. These finding are in line with Bernanke and Gertler (1989), Kiyotaki and Moore (1997), and Wurgler (2000).

Aghion et al. (1999) build a macroeconomic model with micro-foundations, which predicts that less developed financial systems, where credit to the private sector is low, tend to be more volatile. They show that low financial development separates savers from investors, resulting in macroeconomic fluctuations, with the economy converging on a cycle around its steady-state growth path. In contrast, in the presence of financial development, where credit to the private sector is high, the economy converges on a stable growth path along which fluctuations occur due only to exogenous shocks. The authors conclude that, 'when the financial sector is not as well developed, the supply of and the demand for credit is more cyclical.' Thus, 'investors are more likely to get locked out of credit markets when the economy faces a bad shock, only to rush back in when the economy sustains a good shock.'

Table 3: Cross-sectional OLS regression of output volatility for measures of financial development

Financial development	Variable	Output volatility			
		(1)	(2)	(3)	(4)
		2.1	2.2	2.3	2.4
	Growth	-0.233	-0.426**	-0.488***	-0.498***
		(0.184)	(0.163)	(0.126)	(0.163)
	SD (TOT)	0.00306	0.00230	0.00336	0.00282
		(0.00262)	(0.00266)	(0.00269)	(0.00268)
	SD (INF)	-0.000375	-0.000204	-0.000292	-0.000220
		(0.000856)	(0.000860)	(0.000852)	(0.000870)
	Openness	0.508*	0.558*	0.660*	0.517
		(0.383)	(0.388)	(0.387)	(0.391)
Financial depth	PC	-0.614*			
		(0.337)			
Financial efficiency	NIM		0.0550		
			(0.0987)		
Financial stability	ZS			-0.0272*	
				(0.0189)	
Financial access	BB				0.00229
					(0.0109)
	Constant	5.691***	4.822**	5.524***	5.751***
		(1.507)	(1.875)	(1.507)	(1.633)
	Observations	78	77	77	78
	R-squared	0.239	0.212	0.231	0.205
	Adj. R-squared	0.187	0.16	0.18	0.15
	F-stat	7.39	5.39	5.97	2.90
	Shapiro–Wilk test	0.96170	0.9532	0.8715	0.8910
	ovtest	0.2310	0.1921	0.3721	0.0925

Note: Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

The coefficients of net interest margin (column 2) and the number of bank branches per 100,000 adults (column 4) are positive, showing that greater financial access and financial efficiency increase output fluctuations, as found by Popov (2011) and Acemoglu et al. (2003). These findings suggest that financial development plays a mixed role in output volatility, supporting Beck et al. (2006), who conclude that no robust relationship exists between financial development and aggregate economic volatility. The Shapiro–Wilk test statistic and ovtest confirm the normality and functional form of the model.

We find that, as growth increases, output volatility decreases: this result is significant and the signs are correct in all the regressions. When the growth rate of a country increases, this helps stabilize different sectors of the economy and lower the overall output volatility. Since the sources of real and monetary volatility affect producers and intermediaries, we use the SD of TOT changes, *SD(TOT)*, as a proxy for the extent to which an economy is exposed to real sector shocks, and the SD of the inflation rate, *SD(INF)*, as a proxy for the extent to which an economy is exposed to monetary shocks. *SD(INF)* has a negative sign, showing that disturbances or destabilization in the monetary sector do not increase growth volatility. *SD(TOT)* has a positive sign, showing that disturbances or destabilization in the real sector increase output volatility. The signs of the real sector volatility variable are correct and significant. The openness variable has a positive sign in all the regressions, implying that greater openness leads to higher output volatility: economies that are more open are also more vulnerable to global shocks.

The literature on openness and volatility indicates that greater openness can have positive and negative output volatility consequences. For example, Haddad et al. (2013) observe that, 'despite an extensive existing literature on the subject, there is no clear consensus to date whether greater openness comes at the cost of a more volatile growth path.' They argue that, on the one hand, greater openness causes more volatility because an economy that relies on trade to drive economic activity is more exposed to external shocks and thus more volatile. On the other hand, greater openness also reduces volatility because increasing access to international markets protects the economy against significant growth slowdowns due to domestic demand shortages, thereby reducing volatility. Our results are consistent with Kose et al. (2006), who argue that 'openness increases vulnerability to external shocks.' They also observe that 'most studies find that an increase in the degree of trade openness leads to higher output volatility.'

4.2. Cross-Sectional Results with Interaction Terms

Incorporating the interaction terms yields equations 3.1–3.4 (Section 3). Table 4 incorporates two sets of interaction terms, combining financial development with monetary sector volatility and real sector volatility in each regression line. The results in columns 1 and 4 show that the coefficients of private credit to GDP and the number of bank branches per 100,000 adults are positive, which means that an increase in financial depth and financial access increases output volatility. These coefficients are very small and insignificant. Private credit to GDP measures the size of the financial sector: as this sector grows, it attracts greater resources, allowing riskier investments, which can increase output volatility.

Table 4: Cross-sectional OLS regression results with interaction terms

Financial development	Variable	Output volatility			
		(1)	(2)	(3)	(4)
		3.1	3.2	3.3	3.4
	Growth	-0.396**	-0.474***	-0.459***	-0.504***
		(0.177)	(0.167)	(0.131)	(0.165)
	SD (TOT)	0.0115*	-0.00434	0.00933*	0.00295
		(0.00614)	(0.00576)	(0.00603)	(0.00527)
	SD (INF)	0.000217	-0.00320	-0.00367	-0.00107
		(0.00366)	(0.00653)	(0.0141)	(0.00153)
	Openness	0.740*	0.645*	0.683*	0.559*
		(0.413)	(0.395)	(0.391)	(0.404)
Financial depth	PC	0.00601			
		(0.0110)			
	PC-SD (TOT)	-0.000220*			
		(0.000136)			
	PC-SD (INF)	-0.000304			
		(0.000176)			
Financial efficiency	NIM		-0.0781		
			(0.140)		
	NIM-SD (TOT)		0.00144*		
			(0.00115)		
	NIM-SD (INF)		0.000553		
			(0.00120)		
Financial stability	ZS			-0.0120	
				(0.0284)	
	ZS-SD (TOT)			-0.000201*	
				(0.000179)	
	ZS-SD (INF)			0.000290	
				(0.00124)	
Financial access	BB				0.00300
					(0.0169)
	BB-SD (TOT)				-0.0000266
					(0.000245)
	BB-SD (INF)				0.0000679
					(0.0000991)
	Constant	3.786**	5.391***	4.786***	5.637***
		(1.859)	(1.922)	(1.657)	(1.768)
	Observations	78	77	77	78
	R-squared	0.240	0.233	0.247	0.210
	F-stat	4.31	4.22	4.49	2.18
	Shapiro–Wilk test	0.9001	0.9732	0.7501	0.8920
	ovtest	0.2010	0.1561	0.0321	0.5306

Note: Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

An important determinant of the magnitude of the financial sector is the extent of integration with global capital markets. Weaknesses in the country's own financial institutions will matter less if firms have easy access to banks abroad. While greater financial depth could, in principle, serve to smooth the country's adjustment to a shock, it could also increase exposure to adverse sources of dynamic reaction and increase output volatility. The coefficients of net interest margin (financial efficiency) and z-score (financial stability) are negative, which means that financial efficiency and financial stability help decrease output volatility. However, this evidence is weak and insignificant.

The coefficients of *SD(TOT)* are positive in columns 1, 3 and 4 as expected: the more the economy is exposed to real shocks, the higher will be its output volatility. The results suggest that TOT and inflation volatility have a small impact on output volatility. The SD of TOT changes is positive and significant at 1 percent in column 1. The coefficients of *SD(INF)* are negative in columns 2–4, contradictory to our expectations: the more an economy is exposed to monetary shocks, the lower will be its output volatility. The results suggest that TOT and inflation volatility have a small effect on output volatility.

The interaction terms involving real sector volatility and financial efficiency, stability and access have negative signs in columns 1, 3 and 4, respectively. These results provide weak evidence for the dampening effect of financial intermediary development on TOT volatility. Overall, we find only weak evidence to show that the impact of TOT volatility falls in the presence of well-developed financial intermediaries. On the other hand, the positive signs of the interaction term incorporating monetary volatility in columns 2–4 provides weak evidence for the magnifying effect of financial development on monetary volatility. The results suggest that the impact of inflation volatility on output volatility increases in the presence of well-developed financial intermediaries. The Shapiro–Wilk test statistic and ovtest confirm the normality and functional form of the model.

4.3. Two-Stage Least Squares Results for Cross-Sectional Data

This section presents our empirical results, using 2SLS to overcome the limitations of OLS, which does not address the problem of endogeneity. To use OLS, we must be able to meet the zero-conditional-mean assumption. This assumption is violated in three instances: endogeneity (defined as the simultaneous determination of explanatory variables and the explained variable), omitted variable bias and the measurement error in the

explanatory variables. These problems arise for different reasons, but have a common solution: the use of instrumental variables (IVs).

Tables 5 and 6 give the 2SLS results, with legal origin, capital account openness and initial values used as instruments. The R-squared values increase significantly compared to the OLS results. Financial efficiency (net interest margin) and financial stability (z-score) are negative and significant at the 1 percent level of significance, while financial depth (private credit to GDP) and financial access (number of bank branches per 100,000 adults) contribute positively to increasing output volatility. Barring net interest margin, all other measures of financial development contribute significantly to reducing the volatility of the real sector, but the magnitude of this impact is small. The volatility of the monetary sector is also reduced by all financial development indictors except net interest margin.

Table 5: First-stage least squares results

Variable	PC	NIM	ZS	BB
Initial PC	0.5154642***			
	(0.000)			
Initial NIM		0.0407427		
		(0.478)		
Initial ZS			-0.0586756	
			(0.501)	
Initial BB				-0.0155664
				(0.789)
English	19.10035	-1.031573	-4.294696	4.531393
	(0.156)	(0.486)	(0.556)	(0.707)
Socialist	34.6761***	-2.59425	-4.647148	-2.940262
	(0.039)	(0.151)	(0.597)	(0.838)
French	12.28926	-1.2737	-2.278271	2.911096
	(0.356)	(0.387)	(0.753)	(0.807)
German	38.30772***	-1.714287	-1.14229	-7.883624
	(0.015)	(0.302)	(0.889)	(0.562)
Scandan	10.81364	-1.683834	1.867108	-11.77449
	(0.458)	(0.308)	(0.819)	(0.385)
_Cons	-26.87766	8.728035***	18.25015*	12.00845
	(0.196)	(0.000)	(0.073)	(0.499)
Observations	77	73	74	69
R-squared	0.8599	0.7601	0.6089	0.7912
Adj. R-squared	0.84	0.74	0.59	0.76
F-stat	32.74	15.84	7.92	17.68
Sargan's test	(0.6952)	(0.6011)	(0.6689)	(0.4311)

Note: BB = bank branches, NIM = net interest margin, PC = private credit to GDP, ZS = z-score.
Source: Authors' calculations

Table 6: Second-stage least squares results

Financial development	Variable	Output volatility			
		(1)	(2)	(3)	(4)
	Growth	-0.0911	-0.140**	-0.119*	-0.0188
		(0.0601)	(0.0626)	(0.0607)	(0.0667)
	SD (TOT)	0.0986***	0.0129	0.0778**	0.108***
		(0.00742)	(0.0148)	(0.0349)	(0.00847)
	SD (INF)	0.0133*	-0.0569***	0.0892***	-0.000673
		(0.00729)	(0.0127)	(0.0270)	(0.00542)
	Openness	-0.0297	-0.0494	-0.138	0.0744
		(0.156)	(0.154)	(0.167)	(0.158)
Financial depth	PC	0.0430***			
		(0.00811)			
	PC-SD (TOT)	-0.000753***			
		(0.000180)			
	PC-SD (INF)	-0.000634***			
		(0.000187)			
Financial efficiency	NIM		-0.759***		
			(0.162)		
	NIM-SD (TOT)		0.0113***		
			(0.00289)		
	NIM-SD (INF)		0.0114***		
			(0.00290)		
Financial stability	ZS			-0.479***	
				(0.176)	
	ZS-SD (TOT)			-0.00263	
				(0.00246)	
	ZS-SD (INF)			-0.00337**	
				(0.00147)	
Financial access	BB				0.0890***
					(0.0292)
	BB-SD (TOT)				-0.00218***
					(0.000409)
	BB-SD (INF)				-0.000219
					(0.000240)
	Constant	-1.349*	4.424***	10.03***	-2.185**
		(0.754)	(0.995)	(2.735)	(0.834)
	Observations	78	78	78	78
	R-squared	0.884	0.886	0.196	0.885
	Adj. R-squared	0.861	0.861	0.180	0.861
	Wu–Hausman	0.6093	0.5342	0.6661	0.3211
	Cragg–Donald test	55.44	7.42	41.33	85.01
		(0.000)	(0.000)	(0.000)	(0.000
	Cumby–Huizinga test	105.489	105.151	105.002	105.311
		(0.000)	(0.000)	(0.000)	(0.000)
	J-statistic	5.71	6.64	4.57	4.49
		0.33	0.24	0.47	0.48

Note: BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

We have also examined endogeneity for the inclusion of GDP and its variation in the same model, using the Wu–Hausman test, the null hypothesis of which is that the variable under consideration can be treated as exogenous. Here, the test statistics are highly insignificant across all the models, so we cannot reject the null hypothesis of exogeneity. Accordingly, we continue to treat GDP and its variations as exogenous.

4.4. Pooled OLS Results

Table 7 gives the results of the pooled OLS regressions. The negative signs of private credit to GDP, the z-score and number of bank branches show that better financial development helps decrease output volatility. Private credit allows smoother consumption by relieving household liquidity constraints. In turn, smaller fluctuations in real consumption expenditure reduce the volatility of output growth. This indicator has been used in many studies, including Levine and Zervos (1998) and Beck et al. (2000).

Financial intermediaries and markets lower the cost of assessing investments and projects, managing risks and mobilizing savings. All the measures of financial development are significant, although their magnitudes are small. Private credit to GDP and the z-score are significant at 5 percent, while net interest margin and the number of bank branches are significant at 1 percent. When financial depth increases by 1 percent, this causes output volatility to fall by 0.6 percent. An increase in financial efficiency of 1 percent will decrease output fluctuations by 5 percent. When financial stability increases by 1 percent, output volatility falls by 1 percent and an increase in financial access decreases output volatility by 0.9 percent.

Table 7: Pooled OLS results without interaction terms

Financial development	Variable	Output volatility			
		(1)	(2)	(3)	(4)
	Growth	-0.179*	-0.0590	-0.121**	-0.0176
		(0.0917)	(0.0750)	(0.0607)	(0.0967)
	SD (TOT)	0.00485*	0.00623**	0.00677**	0.00493*
		(0.00380)	(0.00281)	(0.00280)	(0.00312)
	SD (INF)	0.000253	0.0676**	0.0614*	0.0410
		(0.000355)	(0.0317)	(0.0317)	(0.0482)
	Openness	0.192	0.603***	0.685***	0.635***
		(0.208)	(0.193)	(0.195)	(0.244)
Depth	PC	-0.00661**			
		(0.00330)			
Efficiency	NIM		0.0502*		
			(0.0428)		
Stability	ZS			-0.0169**	
				(0.00846)	
Access	BB				-0.00919*
					(0.00597)
	Constant	3.513***	-0.408	0.246	-0.377
		(0.909)	(0.987)	(0.842)	(1.162)
	Observations	523	302	303	202
	R-squared	0.053	0.083	0.091	0.069
	VIF	4.18	2.93	2.17	2.32
	Skewness test	0.0021	0.0000	0.0100	0.2781
	Wald test	0.0000	0.0000	0.0000	0.0000
	linktest	0.2577	0.1106	0.1328	0.9677
	ovtest	0.691	0.305	0.420	0.791
	Wooldridge's test	0.603	0.065	0.285	0.000

Note: Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

The results also show that a 1 percent increase in monetary sector volatility raises output volatility from 0.02 percent (column 1) to 7 percent (column 2). As expected, and as Beck et al. (2006) find, *SD(TOT)* has a positive sign, showing that disturbances or destabilization in the real sector will increase output volatility. The results show that a 1 percent increase in real sector volatility raises output volatility from 0.4 percent (column 1) to 0.7 percent (column 3). The coefficient of trade openness has a positive sign in all the regressions, implying that greater openness increases output volatility. An open economy is more exposed to external (global) shocks, thereby increasing its overall output volatility.

The linktest and ovtest results gauge the functional form of the models, based on the idea that, if a regression or regression-like equation is specified correctly, the models are too. The problem of multicollinearity does not arise, as the VIF is less than 10 for all the models. We use the modified Wald test to determine heteroskedasticity, which is resolved by using robust regressions. Since serial correlation in linear panel data models can lead to bias in the standard errors, yielding less efficient results, we apply Wooldridge's test (H0 = no first-order autocorrelation, if $p > 0.1$ accept H0). The results indicate no first-order autocorrelation in models 1 and 3, and weak evidence of first-order autocorrelation in models 2 and 4.

4.5. Pooled OLS Results with Interaction Terms

Table 8 gives the results of the pooled OLS regressions including the interaction terms combining financial development with monetary sector volatility and real sector volatility. The results show that a 1 percent increase in monetary sector volatility will raise output volatility from 0.1 percent (column 1) to 9 percent (column 2). The signs of real sector volatility are correct and significant. The SD of TOT changes is positive in all OLS regressions. The results show that a 1 percent increase in real sector volatility will increase output volatility from 0.3 percent (column 1) to 0.9 percent (column 3). The negative signs of private credit to GDP, z-score and bank branches show that better financial development helps decrease output volatility, which is consistent with Denizer et al. (2002) and Gavin and Hausmann (1996).

Table 8: Pooled OLS results with interaction terms

Financial development	Variable	Output volatility			
		(1)	(2)	(3)	(4)
		3.1	3.2	3.3	3.4
	Growth	-0.183*	-0.0389	-0.119*	-0.0316
		(0.0988)	(0.0735)	(0.0607)	(0.0937)
	SD (TOT)	0.00368	0.00781	0.00924*	0.00465
		(0.00442)	(0.00629)	(0.00689)	(0.00493)
	SD (INF)	0.00140***	0.0953***	0.0931***	0.0102
		(0.000234)	(0.0302)	(0.0358)	(0.0642)
	Openness	0.200	0.591***	0.691***	0.617***
		(0.186)	(0.192)	(0.197)	(0.228)
Financial depth	PC	-0.00694**			
		(0.00316)			
	PC-SD (TOT)	0.00295			
		(0.00696)			
	PC-SD (INF)	-0.000125***			
		(0.000258)			
Financial efficiency	NIM		0.111*		
			(0.0695)		
	NIM-SD (TOT)		-0.000349		
			(0.00139)		
	NIM-SD (INF)		-0.00921*		
			(0.00537)		
Financial stability	ZS			-0.00664	
				(0.0102)	
	ZS-SD (TOT)			-0.000111	
				(0.000183)	
	ZS-SD (INF)			-0.00240**	
				(0.00120)	
Financial access	BB				-0.0130**
					(0.00626)
	BB-SD (TOT)				-0.00331
					(0.00171)
	BB-SD (INF)				0.00393
					(0.00356)
	Constant	3.537***	-0.726	0.0342	-0.135
		(1.067)	(1.010)	(0.831)	(1.153)
	Observations	523	302	303	202
	R-squared	0.155	0.091	0.196	0.173
	VIF	3.53	2.76	2.44	5.37
	Skewness test	0.0150	0.0000	0.0000	0.4258
	Wald test	0.0000	0.0000	0.0000	0.0000
	linktest	0.0000	0.0000	0.0000	0.0000
	ovtest	0.5211	0.1106	0.1328	0.9677
	Wooldridge's test	0.5321	0.3201	0.0632	0.1435

Note: Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

The interaction term *PC-SD(INF)* combines monetary sector volatility with financial depth (private credit to GDP). Its negative sign in column 1 indicates that financial development helps the monetary sector absorb its volatility. We can assume that financial depth does not aggravate monetary volatility and thus decreases overall output volatility. *PC-SD(TOT)* combines real sector volatility with financial development (private credit to GDP). Its positive sign in column 1 indicates that financial intermediaries magnify real sector volatility, in turn implying that financial development increases volatility in the real sector and thus overall output volatility. However, the evidence is very weak and insignificant.

When net interest margin is used as an indicator of financial development, output volatility is positively correlated with the latter in column 2. While the large number of participants in the financial sector allow it to absorb risk and financial markets to become deeper, the broad participation enables risks to be more widely spread across the economy. As Rajan (2006) argues, changes in the financial sector have changed managerial incentives, which in turn have altered the nature of risks undertaken by the system, with some potential for distortion, thereby increasing output volatility.

The interaction terms in column 2 become negative when the financial development indicator changes from private credit to net interest margin. Their negative sign indicates that financial intermediaries dampen monetary and real sector volatility. Based on this, we can argue that financial development complements the financial and real sectors: it does not aggravate volatility in these sectors and helps decrease overall output volatility. The interaction terms in column 3 become negative when the financial development indicator changes to the z-score. Their negative sign implies that financial intermediaries check monetary and real sector volatility. Again, this would indicate that financial development does not aggravate volatility in these sectors, thereby decreasing overall output volatility, although the evidence is very weak and insignificant.

4.6. System GMM Results

We resolve the problem of endogeneity by using IVs. Although the most common solution is to apply 2SLS, this method is more appropriate in the absence of heteroskedasticity for panel data. To apply system GMM, we take the lag of the dependent variable and instrumented endogenous variables, with legal origin and capital account openness as exogenous instruments, along with internal instruments. We use Sargan's test to gauge the validity of the IVs.

The results in Table 9 suggest that both TOT volatility and inflation volatility magnify output volatility. The volatility of inflation (*SD-INF*) is significant in column 3 and its coefficients show that an increase in inflation volatility will raise output volatility significantly by nearly 0.41 percent. TOT volatility is significant in column 1 and its coefficients indicate that an increase in TOT volatility will increase output volatility significantly by nearly 6 percent. These results are consistent with our expectations and the literature. The effect of financial intermediary development is negative, which is consistent with most of the literature. The measures of financial depth (private credit to GDP), financial efficiency (net interest margin) and financial access (number of bank branches per 100,000 adults) are negative and significant.

Table 9: GMM results

Financial development	Variable	Output volatility			
		(1)	(2)	(3)	(4)
		3.1	3..2	3.3	3.4
	L.OV	0.261*	0.108	-0.3267	-0.350*
		(0.213)	(0.221)	(0.226)	(0.150)
	Growth	0.141	-0.0440	0.310*	0.224
		(0.217)	(0.271)	(0.154)	(0.183)
	SD (INF)	-0.000466	-0.231	0.419*	-0.0995
		(0.00656)	(0.231)	(0.295)	(0.109)
	SD (TOT)	-0.0696*	0.0111	0.0115	0.00348
		(0.0408)	(0.0231)	(0.0105)	(0.00746)
	Openness	-1.528*	-0.496	-1.141	-1.654*
		(0.896)	(1.014)	(1.138)	(1.119)
Financial depth	PC	-0.735*			
		(0.411)			
	PC-SD (TOT)	0.0235*			
		(0.0123)			
	PC-SD (INF)	0.000152			
		(0.00229)			
Financial efficiency	NIM		-0.0971		
			(0.269)		
	NIM-SD (TOT)		-0.000242		
			(0.00520)		
	NIM-SD (INF)		0.0413		
			(0.0565)		
Financial stability	ZS			0.0464	
				(0.0365)	
	ZS-SD (TOT)			-0.000149	
				(0.000518)	
	ZS-SD (INF)			-0.0152*	
				(0.00859)	
Financial access	BB				-0.0511**
					(0.0233)
	BB-SD (TOT)				0.00401
					(0.000346)
	BB-SD (INF)				0.0273***
					(0.00978)
	Constant	9.397**	4.611	3.422	7.217
		(3.655)	(3.724)	(4.172)	(4.909)
	Observations	514	296	223	198
	Countries	76	75	75	69
	AR (1)	0.004	0.048	0.068	0.071
	AR (2)	0.969	0.498	0.350	0.323
	Sargan's test	0.909	0.885	0.609	0.562
	Overid test	0.20	0.25	0.15	0.12
	Time dummies	Yes	Yes	Yes	Yes

Note: Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

These results indicate that financial development helps mitigate output volatility. We find weak evidence to support the dampening effect of financial intermediary development on TOT volatility, given that the interaction terms yield mixed signs and insignificant coefficients ranging from 0.01 (column 3) to 2 percent (column 1). There is somewhat stronger evidence to show that financial development magnifies inflation volatility: the results in columns 1, 2 and 4 are positive, with coefficients ranging from 0.018 (column 1) to 4 percent (column 2). The relationship between financial development and output volatility appears to be negative, with real and monetary volatility positively correlated with output volatility. Overall, these results provide only weak evidence that, for a given country, the impact of TOT volatility and inflation volatility diminishes as it develops better financial intermediaries.

For the GMM estimator to be consistent, both instruments chosen from among the lagged endogenous and explanatory variables as well as the assumption that the error terms do not exhibit serial correlation must be valid. To address these issues, we employ Sargan's test, as suggested by Arellano and Bond (1991), Arellano and Bover (1995), and Blundell and Bond (1998). The hypothesis being tested is that the IVs are uncorrelated with a given set of residuals and are therefore acceptable instruments. Since the p-values are greater than 0.1, we cannot reject the null hypothesis and thus conclude that all the instruments are sound.

The purpose of the Arellano–Bond autocorrelation test is to gauge if the error term in levels is not autocorrelated. The values reported for AR (1) and AR (2) are the p-values for first-order and second-order autocorrelated disturbances in the first-differenced equation. Since the p-values are greater than 0.1 for AR (2), we cannot reject the null hypothesis and conclude that all the error terms are uncorrelated. The issue of weak instruments is important to explore because asymptotic theory may be a poor guide to actual finite sample distributions in the presence of weak instruments. In the case of a finite sample, IV estimators may be biased even when they are consistent.

Three issues arise in relation to finite sample bias (Davidson & MacKinnon, 2004). First, when the number of instruments is very large compared to the sample size, the IV estimators may approach OLS estimators and create a similar bias. While this is not important in cross-section analyses, it matters in the case of panel data IV estimators such as the Arellano–Bond estimator. To address this, we restrict the number of instruments. In the case of instrument proliferation, the over-identification

test values approach 1. In our estimations, these values are less than 1 (Table 10). Second, when the correlation between the structural equation error term u and some components of the vector v (first-stage equation errors) is high, then asymptotic theory may be a poor guide to finite sample distributions. Third, in the case of weak instruments, first-stage regressions have a poor fit, which also implies that asymptotic theory is likely a poor guide to the finite sample distribution of the IV estimator.

We follow different approaches to investigate the weak IV problem. The F-statistic shows that the chosen instruments are jointly significant (Table 6). Staiger and Stock (1997) argue that an F-statistic of less than 10 indicates weak instruments. However, this criterion may not be conservative enough when there are many over-identification restrictions. The high values of the R-squared term in the case of the IV estimations (Table 6) indicate that the instruments are not weak. The p-values (Sargan's test) and J-statistics indicate that the instruments are jointly significant and valid. Finally, the p-values generated by Cragg and Donald's (1993) F-statistics also support the validity of the instruments.

4.7. Pooled OLS Results with Dummy and Interaction Term

Table 10 reports the results of the pooled OLS regression when we include the interaction term combining financial development with a dummy for developed countries. The dummy helps gauge whether the role of financial development differs between developed and developing countries in reducing or magnifying output volatility, or whether it has no significant effect at all.

Table 10: Pooled OLS results with interaction of financial development indicators

Financial development	Variable	Output volatility			
		(1)	(2)	(3)	(4)
	Growth	-0.182**	-0.0412	-0.116*	-0.0373
		(0.0920)	(0.0763)	(0.0621)	(0.0969)
	SD (INF)	0.00485	0.00628**	0.00649**	0.00526*
		(0.00380)	(0.00283)	(0.00282)	(0.00313)
	SD (TOT)	0.000238	0.0656**	0.0594*	0.0371
		(0.000355)	(0.0318)	(0.0318)	(0.0481)
	Openness	0.162	0.556***	0.686***	0.580**
		(0.210)	(0.197)	(0.198)	(0.245)
Financial depth	Dummy	-0.282			
		(0.345)			
	PC	-0.00569			
		(0.00392)			
	PC-D	-0.000458			
		(0.00478)			
Financial efficiency	Dummy		-0.447		
			(0.339)		
	NIM		0.0268		
			(0.0482)		
	NIM-D		0.0743		
			(0.0699)		
Financial stability	Dummy			-0.238	
				(0.377)	
	ZS			-0.0188*	
				(0.00966)	
	ZS-D			-0.000987	
				(0.0229)	
Financial access	Dummy				0.492
					(0.338)
	BB				0.00116
					(0.00820)
	BB-D				-0.0182*
					(0.00961)
	Constant	3.702***	-0.208	0.325	-0.202
		(0.922)	(1.001)	(0.846)	(1.173)
	Observations	523	302	303	202
	R-squared	0.056	0.089	0.096	0.086

Note: BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

In columns 1 and 3, the coefficients of the interaction terms (dummy with private credit to GDP and dummy with z-score, respectively) are negative and statistically insignificant. This implies we cannot accept the hypothesis that private credit to GDP and z-scores have a stronger negative relationship with reducing output volatility in developed countries compared to developing ones. Specifically, for every extra US$1 increase in private credit to GDP (financial depth), we will see a 0.000458-point decrease in output volatility in developed countries over and above any effect we see in developing countries. The results in column 1 show that every US$1 increase in private credit to GDP leads to a $(-0.000458 + -0.00569 = -0.00615)$-point decrease in output volatility in developed countries, whereas, a one-point increase in the z-score (financial stability) yields a $(-0.0188 + -0.000987 = -0.01979)$ -point decrease in output volatility in developed countries.

In column 2, the coefficient of the interaction term is positive and statistically insignificant, indicating that we cannot accept the hypothesis that net interest margin as a measure of financial efficiency has a strong positive relationship with magnifying output volatility in developed countries compared to developing ones. Specifically, for every one-unit increase in net interest margin, we will see 0.0743 -point increase in output volatility in developed countries over and above any effect we see in developing countries. In this case, every one-unit increase in net interest margin leads to a $(0.0743 + 0.0268 = 0.1011)$-point increase in output volatility in developed countries.

In column 4, the coefficient of the interaction term is negative and statistically significant, implying that we can accept the hypothesis that the number of bank branches as a measure of financial access has a stronger negative relationship with magnifying output volatility in developed countries compared to developing ones. Specifically, for every one-unit increase in the number of bank branches per 100,000 adults, we will see a 0.0182-point decrease in output volatility in developed countries over and above any effect we see in developing countries. In this case, every additional branch per 100,000 adults leads to a $(-0.0182 + 0.00116 = -0.01704)$ -point decrease in output volatility in developed countries. These results indicate that financial access is the only significant determinant of financial development leading to a decrease in output volatility when a developed country dummy is introduced. The other measures of financial development do not have a significant effect in this context.

4.8. Random Effects and Fixed Effects Results

Finally, we address the issue of unobserved country-specific effects. Table 11 reports the empirical results of the random effects model. All measures of financial development help mitigate output volatility, except for financial efficiency, which is insignificant. Overall, the results suggest that our baseline findings are not sensitive to random effects.

Table 11: Random effects model

Financial development	Variable	Output volatility			
		(1)	(2)	(3)	(4)
	Growth	-0.208**	-0.0629	-0.122*	-0.0176
		(0.0981)	(0.0792)	(0.0626)	(0.0967)
	SD (TOT)	0.00350	0.00555*	0.00637**	0.00493
		(0.00382)	(0.00285)	(0.00282)	(0.00312)
	SD (INF)	0.000272	0.0658**	0.0610*	0.0410
		(0.000353)	(0.0320)	(0.0319)	(0.0482)
	Openness	0.156	0.591***	0.679***	0.635***
		(0.223)	(0.205)	(0.202)	(0.244)
Financial depth	PC	-0.00546			
		(0.00346)***			
Financial efficiency	NIM		0.0498		
			(0.0448)		
Financial stability	ZS			-0.0169**	
				(0.00846)	
Financial access	BB				-0.00919***
					(0.00597)
	Constant	3.865***	-0.304	0.290	-0.377
		(0.972)	(1.042)	(0.868)	(1.162)
	Observations	523	302	303	202
	R-squared	0.28	0.25	0.27	0.20

Note: BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

A fixed effects model is used to address country heterogeneity by controlling for country-specific effects. Table 12 replicates the baseline model with fixed effects. The first column shows that the effect of financial depth is positive and significant, implying that it increases output volatility. The other measures of financial development have no significant impact on output volatility. Overall, the role of financial development is mixed and the variable is sensitive to alternative estimators.

Table 12: Fixed effects model

Financial development	Variable	Output volatility			
		(1)	(2)	(3)	(4)
	Growth	-1.193**	0.163	0.177	-1.387
		(0.579)	(0.849)	(0.855)	(1.575)
	SD (TOT)	-0.00104	0.000239	0.000146	-0.000144
		(0.00445)	(0.00438)	(0.00432)	(0.00735)
	SD (INF)	0.000330	0.0534	0.0550	-0.0140
		(0.000365)	(0.0401)	(0.0399)	(0.0710)
	Openness	-0.427	-0.202	-0.0514	0.474
		(0.534)	(0.755)	(0.755)	(1.290)
Financial depth	PC	0.00965*			
		(0.00568)			
Financial efficiency	NIM		0.0132		
			(0.0948)		
Financial stability	ZS			-0.00646	
				(0.0284)	
Financial access	BB				-0.00274
					(0.0329)
	Constant	13.62***	1.459	0.860	12.25
		(4.225)	(7.109)	(7.153)	(13.90)
	Observations	523	302	303	202
	R-squared	0.15	0.55	0.02	0.03

Note: BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

Tables 13 and 14 summarize the overall effects of financial development on output volatility comparing the cross-sectional, pooled OLS and GMM results.

Table 13: Summary of results without interaction terms

Variable	Cross-section	Pooled OLS	Overall
PC	–	–	–
NIM	+	+	+
ZS	–	–	–
BB	+	–	Mixed
SD (TOT)	+	+	+
SD (INF)	+	+	+

Note: BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

Table 14: Summary of results with interaction terms

Variable	Cross-section	Pooled OLS	GMM	2SLS	Overall
PC	+	–	–	+	Mixed
NIM	–	+	–	–	–
ZS	–	–	+	–	–
BB	+	–	–	+	Mixed
SD (TOT)	+	+	+	+	+
SD (INF)	–	+	–	(+, –)	–
PC-SD (TOT)	–	+	+	–	Mixed
NIM-SD (TOT)	+	–	–	+	Mixed
ZS-SD (TOT)	–	–	–	–	–
BB-SD (TOT)	–	–	+	–	–
PC-SD (INF)	–	–	+	–	–
NIM-SD (INF)	+	–	+	+	+
ZS-SD (INF)	+	–	–	–	–
BB-SD (INF)	+	+	+	–	+

Note: BB = bank branches, INF = inflation, NIM = net interest margin, PC = private credit to GDP, TOT = terms of trade, ZS = z-score.
Source: Authors' calculations

5. Conclusion and Policy Recommendations

Globally high levels of output volatility have raised serious concerns among economists and policymakers. The problem is not only that output volatility hinders sustained and rapid economic growth, but that it also constrains development in general. Financial development has been a key economic and institutional change: it promotes better credit allocation and eases borrowing constraints, which helps reduce the impact of nonfinancial shocks on aggregate output and investment. Our findings show that financial development can have an increasing or decreasing effect on output volatility, depending on which measure is used. Overall, we find evidence of a negative relationship. Financial depth, financial access and financial stability have negative signs and help mitigate output volatility. The role of financial efficiency and financial access in mitigating output volatility is weaker than that of financial depth.

On including interaction terms, we derive mixed results: financial efficiency and financial stability are negatively related to output volatility, while financial depth and financial access are positively related to output volatility. Financial efficiency itself reduces output volatility, but its interaction with monetary sector volatility has a positive sign, indicating that

financial efficiency does not help reduce monetary sector volatility. It appears to play a mixed role, however, in aggravating real sector volatility.

Financial stability is negatively related to output volatility and also helps mitigate volatility in the real and monetary sectors. This shows that financial stability not only reduces output fluctuations directly, but also helps reduce volatility in both sectors. Here, the financial depth variable generates mixed results when considered directly as well as interactively. On the other hand, the financial access interaction term with real sector volatility shows that access helps mitigate real sector volatility. Its interaction with monetary sector volatility shows that financial access magnifies the volatility of the monetary sector. We could also interpret this interaction in view of recent research indicating the negative impact of inflation on financial sector development (see Boyd et al., 2001). This interpretation would characterize inflation volatility as decreasing the ability of financial intermediaries to absorb shocks and thus its capacity to reduce output volatility.

Overall, our results suggest that financial development helps mitigate output volatility, although the evidence remains weak in some cases. We find that financial stability plays a more prominent role in decreasing output volatility than other measures of financial development. The positive sign of real and monetary sector volatility implies that volatility in both sectors contributes significantly to making output more volatile. Finally, our results do not imply that financial sector policies are irrelevant to volatility. Rather, the financial sector plays an important role in smoothing out output fluctuations. In this context, the ownership structure of the banking system, the integration of domestic financial markets with international capital markets, and the regulatory and supervisory framework for banks, among other determinants, could have an important impact on mitigating output volatility.

While these results are strongly suggestive, they do not provide unconditional proof. We do not find any robust interaction between the financial development indicators and our measures of real and monetary volatility. This could be due to the weakness of the indicators (TOT volatility and inflation volatility) or to the limited relevance of this channel. Our empirical analysis should, therefore, be seen as exploratory rather than as providing definite answers.

The study is bound by certain limitations. First, it is inappropriate to generalize from only two proxies for real and monetary shocks. Future

research could, therefore, analyze alternative channels through which financial intermediaries affect output volatility and identify other IVs to improve the quality of results. Second, our findings are sensitive to the fixed effects estimator.

Policy measures are important for the development of the financial sector, given that the effect of financial intermediaries' development through bank selection and investment monitoring could help attract more beneficial investments and increase economic growth. Better developed financial sectors also help reduce transaction costs and risk, which, in turn, promotes long-term investment and stabilizes the economy. This study finds that financial stability plays a more prominent role in decreasing output volatility than other measures of financial development. This implies a need to focus on policies that would help stabilize the financial sector. The central bank and national government should actively implement policies that increase the soundness of the financial sector by providing guarantees to financial institutions and improving their structure, ensuring transparency and accountability, and easing the credit constraints to the banking industry.

We find that real sector and monetary sector volatility contribute significantly to output volatility. This implies that policies seeking to curb output volatility must also minimize volatility in the real sector as well as the monetary sector. Effective fiscal and monetary policy play an important role in reducing output losses and minimizing output volatility. However, our results do not imply that policymakers should refrain from pursuing policies that foster financial development. Rather, they underline that financial intermediaries are not a universal solution to controlling volatility and that policy should also focus on other possible linkages to output volatility.

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Appendix

Table A1: Description of variables

Abbreviation	Variable	Definition
GDP	GDP per capita	GDP divided by midyear population. GDP is the sum of gross value added by all resident producers in the economy plus any product taxes and minus any subsidies not included in the value of the products. It is calculated without making deductions for depreciation of fabricated assets or for depletion and degradation of natural resources. Data in constant 2005 US dollars.
OV	GDP per capita growth	Annual percentage growth rate of GDP per capita based on constant local currency. Aggregates are based on constant 2005 US dollars.
	Import value	Import value indexes are the current value of imports (c.i.f.) converted to US dollars and expressed as a percentage of the average for the base period (2000). UNCTAD's import value indexes are reported for most economies. For selected economies for which UNCTAD does not publish data, the import value indexes are derived from import volume indexes (line 73) and corresponding unit value indexes of imports (line 75) in the IMF's International Financial Statistics.
	Export value	Export values are the current value of exports (f.o.b.) converted to US dollars and expressed as a percentage of the average for the base period (2000). UNCTAD's export value indexes are reported for most economies. For selected economies for which UNCTAD does not publish data, the export value indexes are derived from export volume indexes (line 72) and corresponding unit value indexes of exports (line 74) in the IMF's International Financial Statistics.
INF	Consumer price index (2010 = 100)	Consumer price index reflects changes in the cost to the average consumer of acquiring a basket of goods and services that may be fixed or changed at specified intervals, such as yearly. The Laspeyres formula is generally used. Data are period averages.
TO	Trade openness	Imports of goods and services (% of GDP) + exports of goods and services (% of GDP)
TOT	Terms of trade	Export value/import value*100

Abbreviation	Variable	Definition
PC	Ratio of private credit to GDP	The financial resources provided to the private sector by domestic money banks as a share of GDP. Domestic money banks comprise commercial banks and other financial institutions that accept transferable deposits, such as demand deposits.
NIM	Net interest margin	Accounting value of a bank's net interest revenue as a share of its average interest-bearing (total earning) assets (Bankscope).
ZS	Bank's z-score	Captures the probability of default of a country's banking system. Z-scores compare the buffer of a country's banking system (capitalization and returns) with the volatility of those returns (Bankscope).
BB	Number of bank branches	Total number of ATMs for every 100,000 adults in the reporting country. Calculated as (number of ATMs) * 100,000/adult population.

Source: International Financial Statistics (2014) for financial indicators, remaining indicators from World Development Indicators (2014).

Table A2: List of countries

1	Algeria	21	Dominican Rep.	41	Italy	61	Paraguay
2	Argentina	22	Ecuador	42	Jamaica	62	Peru
3	Australia	23	Egypt, Arab Rep.	43	Japan	63	Philippines
4	Austria	24	El Salvador	44	Jordan	64	Portugal
5	Bahamas	25	Ethiopia	45	Kenya	65	Rwanda
6	Barbados	26	Finland	46	Korea, Rep.	66	Senegal
7	Belgium	27	France	47	Luxembourg	67	Singapore
8	Bhutan	28	Gabon	48	Madagascar	68	South Africa
9	Bolivia	29	Gambia	49	Malta	69	Spain
10	Botswana	30	Germany	50	Mauritius	70	Sri Lanka
11	Burkina Faso	31	Ghana	51	Mexico	71	Sudan
12	Burundi	32	Greece	52	Morocco	72	Sweden
13	Cameroon	33	Guatemala	53	Nepal	73	Switzerland
14	Canada	34	Haiti	54	Netherlands	74	Syrian Arab Rep.
15	China	35	Honduras	55	New Zealand	75	Thailand
16	Colombia	36	Iceland	56	Niger	76	Trinidad & Tobago
17	Costa Rica	37	India	57	Nigeria	77	UK
18	Cote d'lvoire	38	Iran, Islamic Rep.	58	Norway	78	US
19	Cyprus	39	Ireland	59	Pakistan	79	Uruguay
20	Denmark	40	Israel	60	Panama		

Figure A1: Financial development and volatility, developing countries

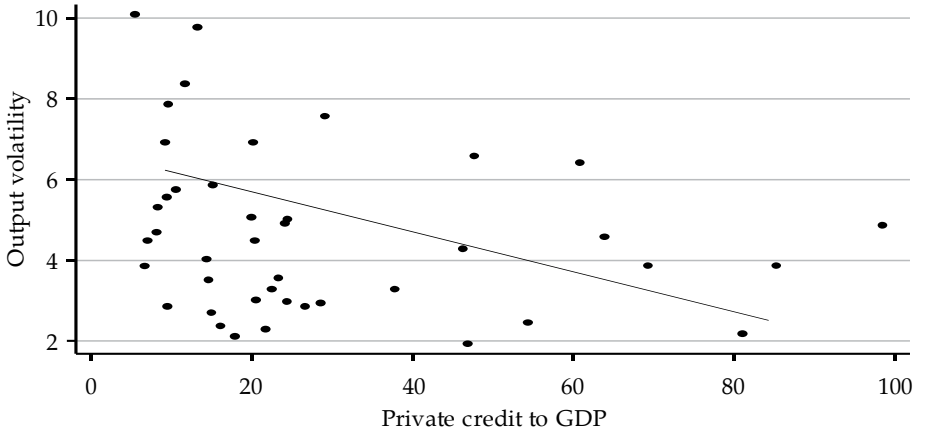


Figure A2: Financial development and volatility, developed countries

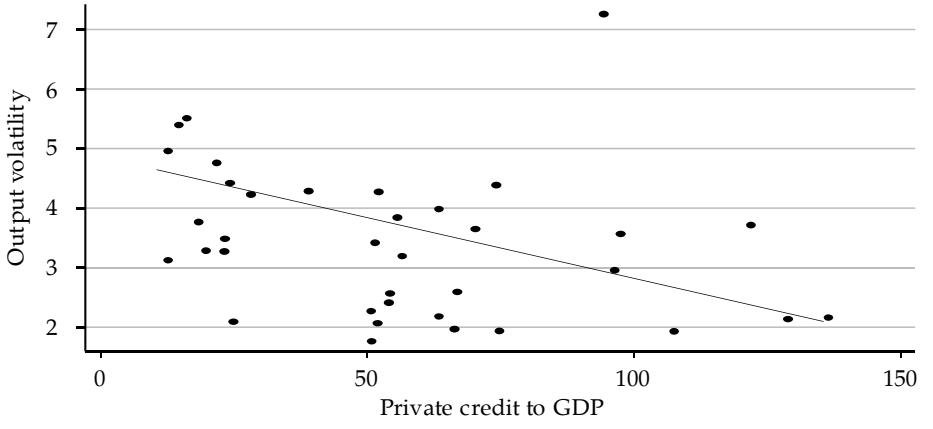


Figure A3: Financial access and output volatility

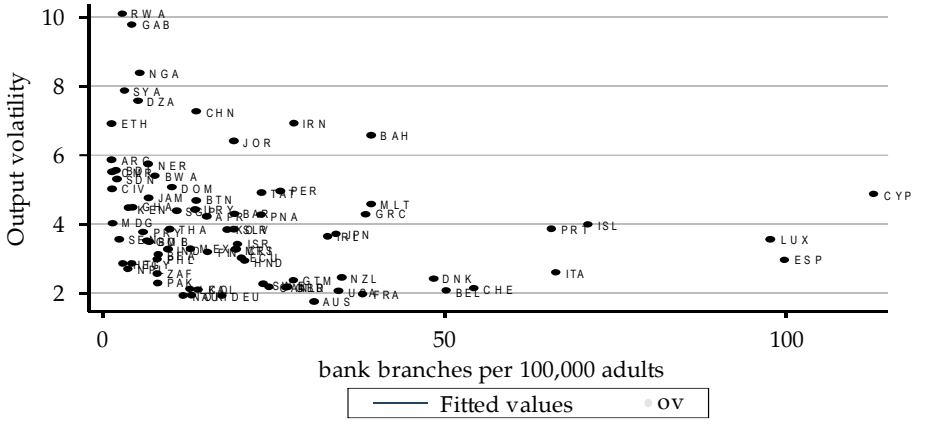


Figure A4: Financial stability and output volatility

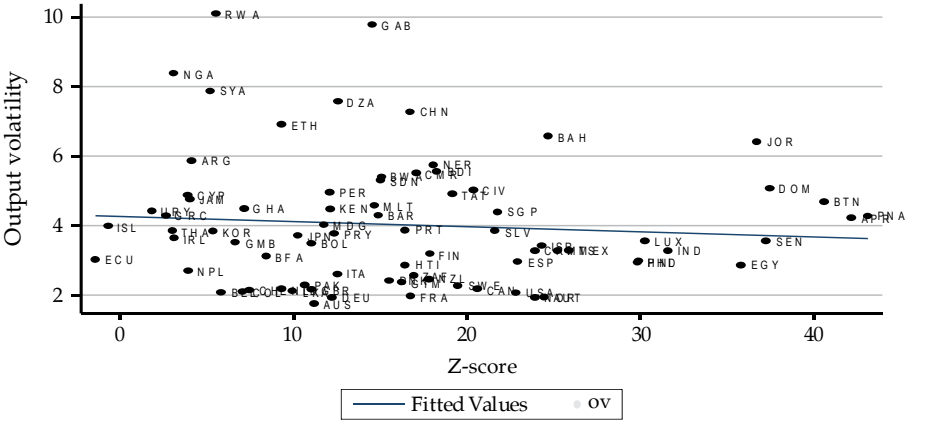


Figure A5: Financial efficiency and output volatility

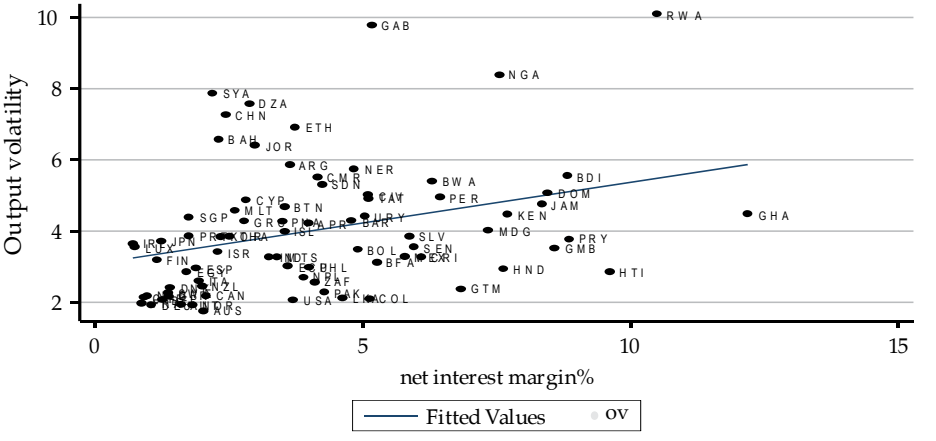
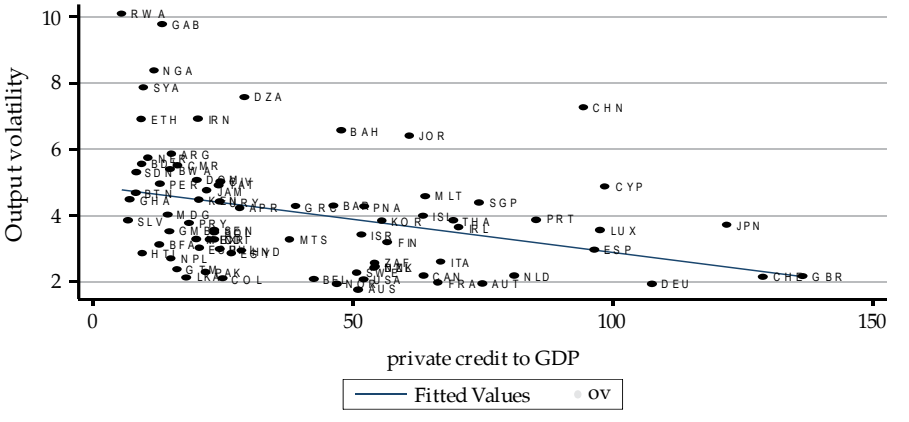


Figure A6: Financial depth and output volatility



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Testing the Governance-Productivity Nexus for Emerging Asian Countries

Ghulam Mustafa* **and Muhammad Jamil****

Abstract

This paper presents panel data estimates of the relationship between governance, aggregate labor productivity (ALP) growth and total factor productivity (TFP) growth for 12 Asian economies between 1996 and 2013. Our results show that government effectiveness has a positive and significant effect on ALP in both levels and first differences. Regulatory quality has a positive impact on ALP only in first difference. Although both government effectiveness and regulatory quality have a positive effect on TFP growth in first difference, only political stability is significant and positive in the levels specification. Other findings indicate that physical capital and human capital have a positive effect on ALP growth. We also find evidence of positive spillover effects with respect to human capital. The positive association between governance, economic growth and productivity provide a better understanding of the role of governance in enhancing economic performance. Our findings have policy implications for ways to achieve good governance to enhance economic growth and productivity.

Keywords: Governance, total factor productivity, economic growth.

JEL classification: C23, D24, E24, O40.

1. Introduction

The Asian economies are expected to overtake the G7 countries in 2018, with China and India being the two largest and fastest growing economies. Consequently, the Asian economies are expected to contribute significantly to future world output growth (Jorgenson & Vu, 2011). However, the quality of governance in Asia is not comparable with that of higher-income countries, nor is it constant across the regional economies. While the rapid growth of the East Asian 'tiger' economies (Singapore, Hong Kong, South Korea and Taiwan) is due to their strong institutional environment, South Asian countries still suffer from weak governance.

* Assistant Professor, Department of Economics, Federal Urdu University of Arts, Science and Technology, Islamabad, Pakistan.
** Assistant Professor, School of Economics, Quaid-i-Azam University, Islamabad, Pakistan.

Since the seminal contribution of North (1991), an increasing number of studies have shown that governance – defined as 'the traditions and institutions by which authority in a country is exercised' (Kaufmann et al., 2009) – plays a significant role in increasing economic growth and development (see Everhart et al., 2009; Kaufmann & Kraay, 2002; Williams & Siddique, 2008). Therefore, 'good governance' has become critical to political as well as economic agendas. Better governance leads to stronger institutions that are better equipped to tackle corruption and enhance the rule of law and political stability in a country. It also helps protect human rights, ensures government accountability to voters and the public, improves the state's capacity to handle external shocks and lowers transaction costs (Ndulu & O'Connell, 1999).

The 'good governance' agenda has engendered various governance reforms to control corruption, improve the judicial system and build political stability. Economists agree that the successful implementation of such reforms would lead to improvements in 'virtually all aspects of the public sector' (Rodrik, 2004; Grindle, 2004). On the other hand, 'bad governance' is recognized as a major impediment to growth and economic development, with many African countries being a case in point (Moore, 2001). Collier (2007) identifies the 'bottom billion people' of the world's population as living in 58 small countries that lag behind other developing countries in terms of growth and poverty reduction. The study shows that 70 percent of the 'bottom billion' live in Africa and that 76 percent of these countries have undergone prolonged 'bad governance'. In sum, 'good governance' is necessary to achieve faster growth and development.

This paper contributes to two strands of the literature: (i) the current debate on the role of governance quality in growth (see Acemoglu et al., 2002, 2004; Bhattacharyya, 2009; North, 1991) and (ii) macroeconomic studies of the relationship between governance and total factor productivity (TFP) (Olson et al., 2000). While the literature shows that governance has a positive impact on economic growth (see Kaufmann & Kraay, 2002; Cebula & Ekstrom, 2009), to our knowledge there is no existing study on the relationship between governance, average labor productivity growth and TFP growth for Asian countries.

Our objective is to examine how different dimensions of governance affect productivity across a sample of Asian countries, using aggregate labor productivity (ALP) and multifactor productivity (MFP) as measures of productivity. ALP growth measures economic development within a country in terms of growth, competitiveness and living standards (OECD,

2001). MFP growth captures improvements in technology due to organizational and institutional changes, changes in returns to scale, the impact of unmeasured inputs such as research and development, other intangible factors and measurement errors, and disembodied technological change (see Erumban, 2008; Inklaar et al., 2008; van Ark et al., 2008).

This paper contributes to the debate on the relationship between governance, growth and productivity at the country level, based on six dimensions of governance. Using ALP and MFP to measure economic performance, we assess the internal returns to governance as well as its potential to generate important spillover effects. The literature suggests that reliable data on governance is a key problem in identifying the valid association between governance and growth (Williams & Siddique, 2008).

Our results indicate that government effectiveness and regulatory quality have positive and significant effects on ALP growth and MFP growth. The results of the ALP specification in levels show that only government effectiveness is statistically significant at conventional levels. The long-run MFP specification indicates that political stability has a positive and significant effect on conventional levels. To determine the robustness of these results, we also account for the impact of different levels of education – primary, secondary and tertiary. Our main findings remain robust to changes in human capital. We also find that governance can generate spillover effects.

The rest of the paper is organized as follows. Section 2 reviews key studies on the impact of governance on economic growth and productivity. Section 3 describes the data used and related issues. Sections 4 and 5 present our empirical strategy and results as well as a robustness analysis. Section 6 concludes the study.

2. Literature Review

Economists have long striven to explain what causes economic growth. The vast literature includes cross-country analyses that trace growth to physical capital, human capital, trade policy, financial development and geography (see Barro, 1991; Dollar & Kraay, 2003; Levine & Renelt, 1992; Rigobon & Rodrik, 2005). However, since the mid-1990s, the quality of institutions and/or governance has been put forward as one of the most important explanations for economic growth. For example, Rodrik et al. (2004) conclude that the quality of institutions 'trumps' everything else. Better governance promotes economic growth via an ordered and more

transparent environment, incentives for capital accumulation and improvements in technology (North, 1991).

This section presents an overview of the main studies on the role of governance in economic growth. Sen (1999) suggests that democracy is 'an essential component of the process of development.' Similarly, Earle and Scott (2010) argue that democracy helps promote harmony among citizens, reducing conflict and enhancing political stability. Halperin et al. (2005) show that the increase in prevalence of democracy reduced armed conflict and civil war between 1991 and 2003. Using data on revolutions, coups and political assassinations as proxies, Barro (1991) includes a measure of political instability in his growth regressions and shows that these measures are inversely related to growth and investment. His findings imply that democracy is beneficial for growth.

Similarly, Rivera-Batiz (2002) develops a general equilibrium framework to show that democracy improves the quality of governance and fosters economic growth. This model shows that strong democratic institutions have a positive impact on growth by restricting the actions of corrupt officials. Democracy thus reduces corruption and raises economic performance by stimulating technological progress. The empirical findings indicate that democracy was an important determinant of TFP growth during 1960–1990 for a cross-section of countries. Studies on the relationship between political regimes and economic growth are, however, divided as to whether democracy benefits or harms growth. For instance, Gerring et al. (2005) suggest that 'democracy has either a negative effect on GDP growth or no overall effect.'

There is considerable evidence to suggest that the quality of governance matters for economic growth, both in developing and developed countries. Hall and Jones (1999) argue that institutional differences are the main cause of cross-country differences in productivity and GDP per capita. Tracing the influence of institutions to the influence of Western European settlements or colonization and their adaptation to social infrastructure (institutions), they find that institutions have promoted productivity and growth. Cavalcanti and Novo (2005) obtain similar estimates, finding that a 1 percent increase in institutions is associated with a 5 percent increase in GDP per worker, based on a cross-sectional analysis of data for 1988.

Acemoglu et al. (2002) and Rodrik et al. (2004) focus on the different incentives for imposing extractive versus economic institutions in different parts of the world, using settler mortality rates as an instrument. In an

influential study, Acemoglu et al. (2002) show that a 'reversal of fortune' took place in former European colonies: economies that were rich in the sixteenth century are, today, poor. Their evidence does not support the geography hypothesis, which traces income differences to differences in geography. The authors argue that European colonists introduced economic institutions in densely settled areas and imposed extractive ones in poor and unsettled areas. This institutional reversal caused a reversal in income levels.

Rodrik et al. (2004) provide econometric evidence to show that the quality of institutions is the most important factor in enhancing income levels. For a sample of 137 countries in a cross-sectional analysis of data for 1995, they find that trade becomes negative and insignificant while geography has only a weak effect once institutions are controlled for. Furthermore, the identity of the colonial power and the percentage of the European population do not influence income per capita.

Corruption – the misuse of public power for private benefits – comes in many forms, affecting efficiency and investment (Bardhan, 1997). Corruption has a negative influence on economic growth because it lowers the private marginal product of capital, for example, by acting as a tax on the proceeds of investment, by causing inefficiencies in the private and public sectors, by distorting the financial process and environment and, to some extent, by causing political instability and anarchy. In causing the misallocation of production between sectors, corruption discourages domestic and foreign investment, innovation and good governance (Bardhan, 1997; Everhart et al., 2009; Mauro, 1995).

Everett et al. (2007) suggest that corruption creates mistrust of the political government, increases the cost of public infrastructure projects, weakens the rule of law and decreases government revenues. However, some studies argue that corruption has a positive effect on growth because it serves to 'grease the wheels' of a 'rigid administration', thereby proving beneficial in slow bureaucratic countries (Huntington, 1968; Leff, 1964; Leys, 1965). Mauro's (1995) cross-country analysis shows that corruption has a negative influence on investment and economic growth. The results indicate that a one-standard deviation improvement in corruption leads to a five-percentage point increase in investment, with annual GDP growth rising by more than half a percentage point. The empirical findings support the stylized fact that poor countries tend to have less efficient institutions than developed countries and that these lags persist and lead to less growth and greater poverty in the future.

Everhart et al. (2009) find further evidence of the negative effect of corruption on economic growth. Their empirical findings indicate that corruption has an unambiguous negative effect on governance, which further deters economic growth. However, the effect of corruption on investment is ambiguous as opposed to earlier findings in the literature: their results indicate a positive but insignificant effect on public investment while the effect of corruption on private investment is insignificant. This suggests that corruption does not affect private investment, but it does so indirectly through public investment. The results show that a one-percentage point increase in corruption is associated with a 0.124 percentage point decline in governance.

Cebula and Ekstrom (2009) investigate different forms of economic freedom and dimensions of governance to gauge their impact on growth in OECD countries between 2004 and 2007, while adjusting for G8 status and budget status. They use five forms of economic freedom, with trade freedom as an interesting addition to earlier investigations. Trade freedom reflects the openness of an economy in importing goods and services, and the free interaction of citizens in international buying and selling. Governance is measured by control of corruption, political stability and perceived corruption. The study's panel least squares estimates suggest that governance has a positive impact and corruption a negative impact on economic growth. Economic freedom – as measured by trade freedom, business freedom, monetary freedom and property rights – has a positive effect on economic growth.

Li and Samsell (2009) suggest that 'governance matters' for trade flows between two trading partner countries. Therefore, rule-based countries trade more than relation-based ones. Their study suggests that rule-based countries attract trade flows, given their transparent regulation and presence of fair rules. However, it is more difficult and/or costlier to trade with relation-based countries due to their poor governance environment.

3. Data and Summary Statistics

This study uses annual panel data for 12 Asian countries over the period 1996–2013. The sample countries include Bangladesh, India, Pakistan, Sri Lanka, Hong Kong, Indonesia, Malaysia, the Philippines, Singapore, South Korea, Thailand and China. Our main source of data is the Conference Board database, which provides information on real GDP (at purchasing power parity exchange rates) and employment for most of the

countries in this sample.[1] This is complemented by the World Development Indicators series for gross fixed capital formation, which is then used in capital stock calculations.[2] World Governance Indicators (WGIs) are obtained from Kaufmann et al. (2011). We estimate the real capital stock series using the perpetual inventory method (PIM), based on the World Development Indicators investment series.

Human capital is an intangible asset and therefore difficult to measure. The school attainment data from Barro and Lee (2010) includes the share of persons who have completed primary, secondary and tertiary education in a sample of people aged 25 and over. This data is available in five-year periods for 1950–2010 for a large set of countries. We compare the performance of two measures – the Cohen and Soto (2007) series and the Barro and Lee (2010) series. The latter is one of the most commonly used measures of human capital in the macro literature on developing countries. In relation to the quality of schooling data, Krueger and Lindahl (2001) use reliability ratios to compare Barro and Lee (2010) and Kyriacou (1991). The authors suggest that the Barro and Lee dataset is more advanced and more reliable than Kyriacou's education estimates. In addition, it uses census information where available, relying on enrolment data and a PIM to fill in the missing values (de la Fuente & Doménech, 2006).

Cohen and Soto (2007) claim greater precision than Barro and Lee, serving as a direct substitute for the latter. Their measure uses information on educational attainment by age group. Accordingly, we use data on the average years of schooling for the population aged 15 and above from Barro and Lee (2010) to construct a Cohen–Soto measure of human capital. Next to aggregate measures of human capital, we look at the effect of primary, secondary and tertiary education on growth. For this sample of Asian countries, where levels of education are, in some cases, still very low compared to developed countries, changes in basic education could have a greater impact on performance.

Reliable data on governance is an important problem in identifying a valid association between governance and growth (Williams & Siddique, 2008). Using the WGI data obtained from Kaufmann et al. (2011), we consider six dimensions of governance: voice and accountability, political stability and absence of violence/terrorism, government effectiveness, regulatory quality, the rule of law and control of corruption (Table 1). This choice is justified, given that the WGIs are related to Transparency

[1] http://www.conference-board.org/data/economydatabase/
[2] Accessed through www.esds.ac.uk

International's corruption perceptions index. The WGIs are better because they cover six dimensions of governance, including corruption, span more countries and are available consistently for our sample of countries. The WGIs are based on the subjective views expressed by numerous private firms, citizens, think-tanks, NGOs and international organizations on the quality of governance in a country. We have relied on these informed stakeholders for the following reasons.

- Perceptions, impressions and views influence the actions of agents. If agents do not trust the judiciary or police, they are unlikely to avail these services. These perceptions are very important for investment decisions by businesses and voting decisions by citizens.
- Subjective measures are the only choice in the case of 'corruption' where we cannot measure the defined indicator by any objective data.
- Objective indicators capture a de jure notion of laws, which is significantly different from the de facto reality on the ground.
- The choice between subjective and objective indicators of governance reflects a false dichotomy: all measures of governance inherit some degree of judgement.

Table 1: Six dimensions of governance

Dimension	Captures
Voice and accountability	Perceptions of the extent to which a country's citizens can participate in selecting their government, as well as freedom of expression, association and the media.
Political stability and absence of violence	Perceptions of the likelihood that the government will be destabilized or overthrown by unconstitutional or violent means, including politically motivated violence and terrorism.
Government effectiveness	Perceptions of the quality of public services, the quality of the civil service and the degree to which it is independent from political pressure, the quality of policy formulation and implementation, and the credibility of the government's commitment to such policies.
Regulatory quality	Perceptions of the ability of the government to formulate and implement sound policies and regulations that permit and promote private sector development.
Rule of law	Perceptions of the extent to which agents have confidence in and abide by the rules of society, particularly the quality of contract enforcement, property rights, the police and the courts as well as the likelihood of crime and violence.
Control of corruption	Perceptions of the extent to which public power is exercised for private gain, including both petty and grand forms of corruption, as well as 'capture' of the state by elites and private interests.

Source: Kaufmann et al. (2009).

The governance data is available biannually from 1996 and annually between 2002 and 2010. All six governance indicators range from –2.5 to 2.5, with greater numerical values representing stronger governance performance (Kaufmann et al., 2009). Table 2 presents descriptive statistics for the variables used, including the governance indicators. The smallest value (–1.681) for voice and accountability (VOACC) is that of China in 2006, while the largest value (0.729) is that of South Korea in 2010. This indicates that Chinese citizens have the least VOACC, while South Korean citizens have the highest VOACC. Pakistan is the least politically stable country, as shown by the political stability and absence of violence (POLST) indicator for 2011, while Singapore in 2008 has the highest POLST score. Bangladesh fares the worst in terms of government effectiveness (GOVEFF), regulatory quality (REGQU), rule of law (RLAW) and control of corruption (CONCORR) in 2005, 2004, 2003 and 2004, respectively. Singapore has the highest score for these four indicators in 2008, 1996, 2012 and 2004.

Table 2: Summary statistics

Variable	Obs.	Mean	SD	Min	Max
Real GDP (Y)	216	1,070,689	2,025,215	56,954.61	12,600,000
Labor (L), in thousands	216	121,640	216,519	1,800	770,000
Physical capital stock, in billions	216	1,250	2,200	32.6	15,200
Average years of schooling (Barro and Lee)	216	7.50	2.21	3.39	12.08
Primary education	216	4.15	1.09	1.78	5.82
Secondary education	216	2.64	1.14	1.05	5.06
Tertiary education	216	0.36	0.31	0.04	1.20
Average years of schooling (Cohen and Soto)	216	4.69	0.83	2.10	5.73
Voice and accountability (VOACC)	216	-0.215	0.607	-1.681	0.729
Political stability and absence of violence (POLST)	216	-0.543	1.018	-2.812	1.343
Government effectiveness (GOVEFF)	216	0.327	0.867	-0.865	2.430
Regulatory quality (REGQU)	216	0.237	0.890	-1.095	2.247
Rule of law (RLAW)	216	0.109	0.799	-1.020	1.770
Control of corruption (CONCORR)	216	0.003	0.997	-1.488	2.417

Note: Y = real GDP (in 1990 US$ million, converted at Geary–Khamis PPP), L = employment (in thousands), K = physical capital stock. The education variables represent average years of schooling attained (primary, secondary and tertiary) by population aged 15 years and over, based on Barro and Lee (2010). Cohen–Soto measure = average years of schooling attained computed using Cohen and Soto's (2007) methodology, based on Barro and Lee's (2010) data on total average years of schooling.
Source: Authors' calculations.

Figures 1 to 6 illustrate the positive correlation between these dimensions of governance and ALP for our sample of countries. While the positive association between the variables of interest is important, this correlation does not imply causal relationships and needs further analysis through a series of econometric models.

Figure 1: The Association Between Log ALP and VOACC

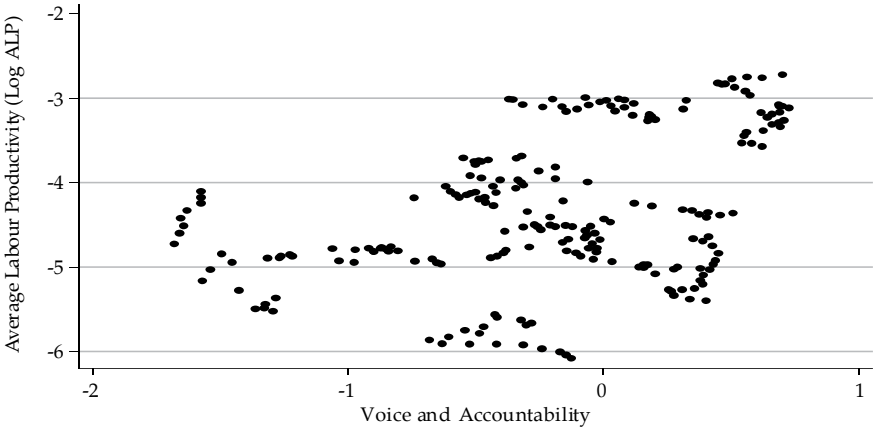


Figure 2: The Association Between Log ALP and POLST

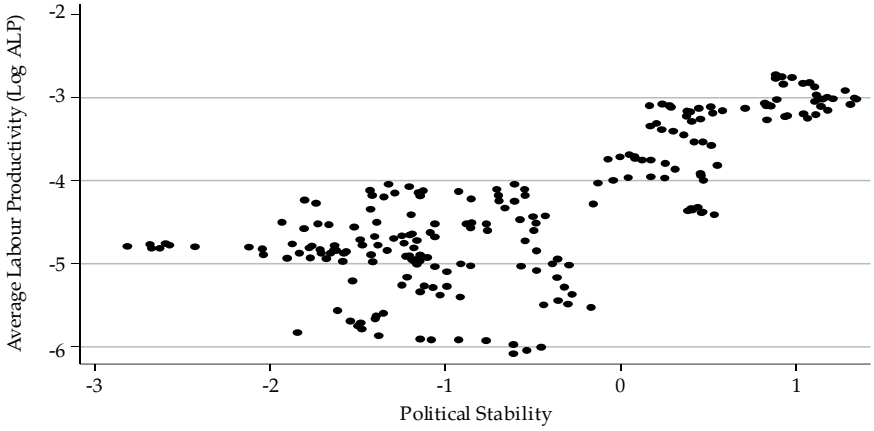


Figure 3: The Association Between Log ALP and GOVEFF

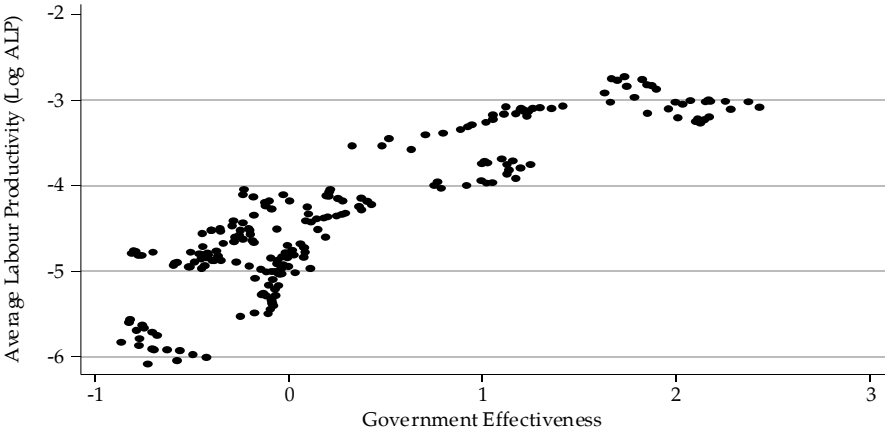


Figure 4: The Association Between Log ALP and REGQUA

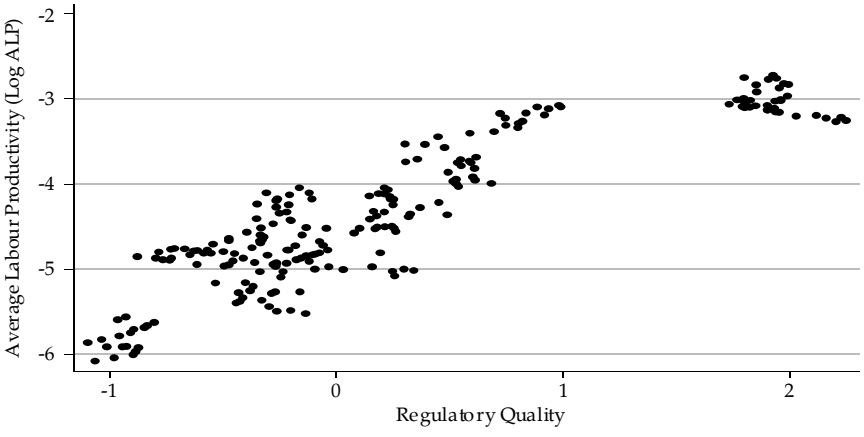


Figure 5: The Association Between Log ALP and RULAW

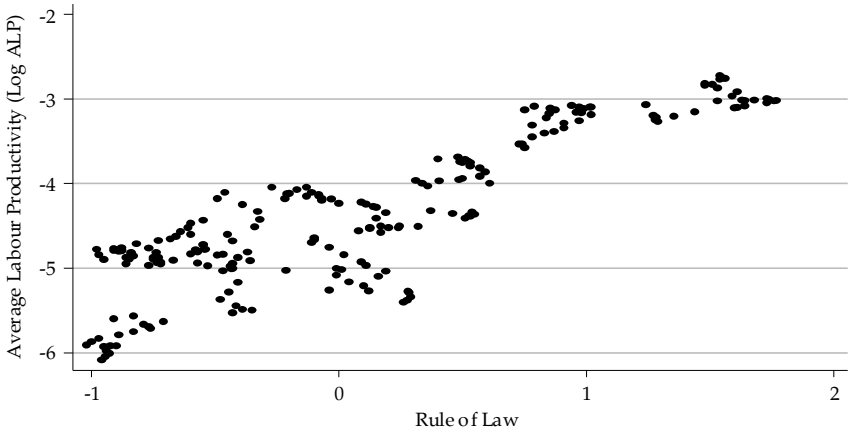
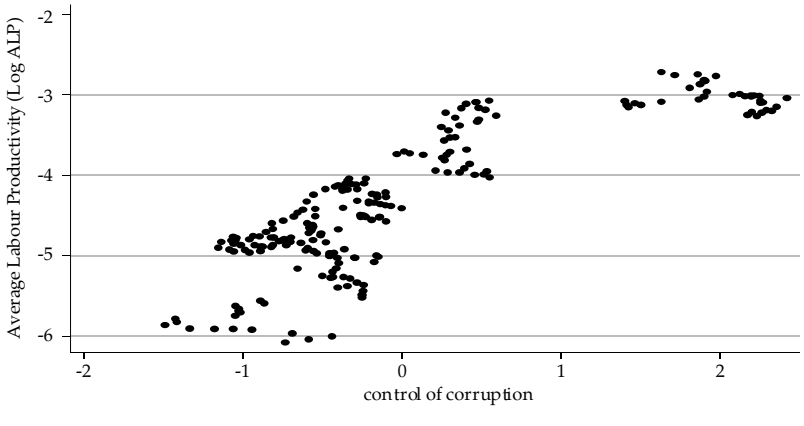


Figure 6: The Association Between Log ALP and CONTCORR



4. Methodology and Empirical Results

We start our analysis of the impact of governance on ALP with an augmented neoclassical production function. This framework models value added (*Y*) as a function of the total number of workers (*L*), aggregate physical capital stock (*K*) and the stock of human capital (*H*).

$$Y_{it} = A_{it}F_i(L_{it}, K_{it}, H_{it}) \tag{1}$$

where $i = 1, 2, 3, \ldots N$ denotes the sampled countries, $t = 1, 2, 3, \ldots T$ denotes the period and A_{it} is MFP. $F_i(\cdot)$ indicates diminishing returns in factor inputs and constant returns to scale in production technology. Under the Cobb–Douglas production function assumptions, we can rewrite the production function (1) in per capita terms as follows:

$$ln\,(Y/L)_{it} = \alpha_0 + \alpha_1 ln(K/L)_{it} + \alpha_2 ln(H/L)_{it} + \varepsilon_{it} \tag{2}$$

An emerging strand of the literature argues that improvements in the quality of institutions increase the accumulation of human capital and contribute to long-run economic growth (Bhattacharyya, 2009; Dias & Tebaldi, 2012). We include the impact of the six governance indicators by extending equation (2) as follows:

$$ln\,(Y/L)_{it} = \emptyset_0 + \emptyset_1 ln\left(\frac{K}{L}\right)_{it} + \emptyset_2 ln\left(\frac{H}{L}\right)_{it} + \emptyset_3(GOV)_{it} + \eta_i + \mu_{it} \tag{3}$$

This long-run growth regression models the log of ALP, $ln(Y/L)_{it}$, as a function of (i) physical capital intensity, $ln(K/L)_{it}$, (ii) the stock of human capital per worker, $ln(H/L)_{it}$ and (iii) governance, GOV_{it}. The term u_{it} denotes other stochastic factors influencing ALP growth. Next, we follow Bhattacharyya (2009) in taking first differences of equation (3) to obtain the following equation:

$$\Delta ln(Y/L)_{it} = \pi_1 \Delta ln\left(\frac{K}{L}\right)_{it} + \pi_2 \Delta ln\left(\frac{H}{L}\right)_{it} + \pi_3 \Delta(GOV)_{it} + (\mu_{it} - \mu_{it-1}) \tag{4}$$

where Δ represents the first difference of the variable in question. The growth model in equation (4) is based on the models used by Benhabib and Spiegel (1994), Temple (1999) and Mankiw et al. (1992). We thus account for the role of the governance indicators by extending the Cobb–Douglas production function framework to include these dimensions, using the WGI data.

We use three measures of human capital: (i) the stock of human capital, drawing on Barro and Lee (2010); (ii) three different levels of education (primary, secondary and tertiary); and (iii) average years of schooling, applying Cohen and Soto's (2007) methodology to the Barro and Lee (2010) data. Since the strong correlation between the six governance indicators (Table 3) leads to their joint inclusion in the general productivity growth equation,[3] we run six separate regressions to evaluate the impact of each indicator separately.

Table 3: Correlation between ALP and governance indicators

	VOACC	POLST	GOVEFF	REGQU	RLAW	CONCORR
VOACC	1.000					
POLST	0.394	1.000				
GOVEFF	0.406	0.877	1.000			
REGQU	0.448	0.832	0.947	1.000		
RLAW	0.565	0.857	0.927	0.925	1.000	
CONCORR	0.418	0.856	0.949	0.963	0.933	1.000

Source: Authors' calculations.

Table 4 gives the results of ordinary least squares (OLS) regressions of ALP growth on the growth of physical capital per worker and human capital per worker, and various dimensions of governance. Our estimation strategy allows us to use heteroskedasticity-consistent standard errors. We therefore model governance as a factor in the ALP growth regressions, with country fixed effects and time dummies to account for endogeneity and cross-country heterogeneity. This strategy allows us to avoid historical instruments and use the more reliable WGIs. Here, we employ the Barro and Lee (2010) measure of human capital.[4] The first-difference specification addresses any nonstationarity problems that might arise, given the long period spanning the panel.

Modeling the relationship between institutions and growth is subject to the problems of endogeneity and reverse causality, as high-income countries with greater productivity levels seem to have better institutions (Rodrik et al., 2004). Consequently, the literature on the long-run impact of institutions on growth relies on historical and geographical instruments.

[3] Since the correlation between VOACC and the other indicators is relatively small, it is used initially in addition to the other governance indicators in all the regressions. However, the VOACC coefficient is never statistically significant and thus these results are not presented here.

[4] Similar results were obtained using the Cohen–Soto measure, although the human capital variable does not remain significant (see Appendix).

Acemoglu et al. (2002) and Rodrik et al. (2004) use settler mortality as an instrument for schooling and institutions. Although this has strong explanatory power in the first-stage regression, it encounters the problem of multicollinearity in the second stage of a 2SLS regression. This implies that such cross-country studies on the relationship between institutions and growth are not as meaningful (Dollar & Kraay, 2003; Glaeser et al., 2004; Bhattacharyya, 2009).

To account for the multicollinearity problem, Bhattacharyya (2009) uses a dynamic panel regression model and the five-year average of the natural logarithm of real GDP per capita as the dependent variable to isolate the partial effects of institutions and human capital on growth. The study does not encounter any multicollinearity problems since the preferred instruments' lagged levels and the lagged differences of institutions are valid under the generalized method of moments. Therefore, we model governance as a factor in the ALP regressions with country fixed effects and time dummies to account for endogeneity and cross-country heterogeneity. Again, this allows us to avoid historical instruments and use the more reliable WGIs instead.

Table 4: Impact of governance on ALP growth, using Barro and Lee measure

	First-difference specification						
	VOACC	**POLST**	**GOVEFF**	**REGQU**	**RLAW**	**CONCORR**	**GOVER**
$\Delta\log(K/L)_t$	0.423***	0.424***	0.438***	0.439***	0.412***	0.412***	0.443***
	(0.114)	(0.115)	(0.115)	(0.114)	(0.115)	(0.115)	(0.112)
$\Delta\log(H/L)_t$	0.280**	0.280**	0.285**	0.292**	0.286**	0.288**	0.272**
	(0.127)	(0.127)	(0.128)	(0.130)	(0.127)	(0.129)	(0.124)
$\Delta\log(GOV)_t$	0.009	0.006	0.046**	0.035*	0.005	0.001	0.034
	(0.015)	(0.009)	(0.021)	(0.020)	(0.020)	(0.017)	(0.030)
R-sq.	0.682	0.682	0.691	0.689	0.681	0.681	0.685
Obs.	204	204	204	204	204	204	204
	Levels specification						
	VOACC	**POLST**	**GOVEFF**	**REGQU**	**RLAW**	**CONCORR**	**GOVER**
$\Delta\log(K/L)_t$	0.644***	0.659***	0.643***	0.649***	0.658***	0.648***	0.648***
	(0.026)	(0.031)	(0.031)	(0.029)	(0.028)	(0.029)	(0.029)
$\Delta\log(H/L)_t$	0.216***	0.142***	0.177***	0.149***	0.154***	0.169***	0.169***
	(0.044)	(0.046)	(0.047)	(0.048)	(0.045)	(0.053)	(0.053)
$\Delta\log(GOV)_t$	-0.059***	-0.012	0.068**	0.025	-0.057***	-0.016	-0.016
	(0.014)	(0.009)	(0.027)	(0.023)	(0.014)	(0.027)	(0.027)
R-sq.	0.998	0.998	0.998	0.998	0.998	0.998	0.998
Obs.	216	216	216	216	216	216	216

Note: *** = significant at 1%, ** = significant at 5%, * = significant at 10%. GOVER is the unweighted governance index derived from the average of the six dimensions of governance. In the first-difference specification, the dependent variable is log ΔALP. In the levels specification, the dependent variable is log ALP. Heteroskedasticity-consistent standard errors given in parentheses. All equations include country and year dummies.
Source: Authors' calculations.

The upper part of Table 4 presents the OLS estimates based on equation (4), using the Barro and Lee (2010) human capital measure for the 1996–2013 period. This is a two-way fixed effects model, as we control for both country and time effects. Therefore, our estimates are unbiased since they are not influenced by omitted time-variant or country-variant factors. Furthermore, the panel data estimates are not affected by cross-sectional dependence because the period in question is not very long. The results of the first-difference specification clearly indicate that governance has a positive impact on ALP growth. The estimated coefficient of GOV is positive throughout, although it is significant for only two indicators – government effectiveness (GOVEFF) and regulatory quality (REGQU). The lack of statistical significance in the case of the other four indicators could be due to the small sample used and problems associated with the first-difference estimation technique.

Adams-Kane and Lim (2014) describe two channels by which government effectiveness and regulatory quality can improve productivity and economic growth. First, an effective government can theoretically implement a sound macroeconomic policy that enhances the economic progress of the country. Second, an effective government leads to better public financial management, with effective spending thus increasing productivity and economic growth. Our findings provide evidence that human capital has a positive and significant impact on ALP growth in first differences as well as levels. These results are consistent with previous studies showing that human capital has a positive impact on long-run economic growth (see Barro, 1991; Mankiw et al., 1992; Temple, 1999). We find that physical capital per worker has a strong positive effect on ALP growth. Overall, our results underscore the significance of governance, suggesting that government effectiveness and regulatory quality are important factors in increasing ALP in this sample of Asian economies.

The analysis so far suggests that governance has a positive effect on economic ALP growth in these economies. However, it is well understood that institutional variables change slowly over time. Accordingly, we adopt a long-run specification. Table 4 also gives the fixed effects regressions of ALP in levels. These are based on equation (3), including controls for country effects and time dummies. In these regressions, the dependent variable is ALP, with physical capital per worker, human capital per worker and governance as the independent variables. These results are the long-run estimates of the impact of governance on economic performance.

In general, the results of the levels specification in Table 4 are robust to the first-difference results. In the long-run specification, only government effectiveness (GOVEFF) has a positive and significant effect on ALP, as before. However, VOACC does not remain positive and significant in the results given in Table 5, implying that changes in government structure are more relevant for ALP. Human capital per worker has a positive and significant impact on ALP. Similarly, the coefficient of physical capital per worker remains significant. Overall, the results imply that governance has a positive effect on ALP in this sample of Asian economies.

Next, we examine the impact of the dimensions of governance on MFP growth. Not only does a country's institutional framework have a direct impact on ALP growth, but it can also affect the country's ability to enjoy the spillover effects of governance. Accordingly, we specify equation (5) for MFP growth, following Mason et al. (2012) and Vandenbussche et al. (2006):

$$\Delta MFP_{it} = \varphi_1 \Delta \ln(K/L)_{it} + \varphi_2 \Delta \ln(H/L)_{it} + \varphi_2 \Delta(GOV)_{it} + (\mu_{it} - \mu_{it-1}) \quad (5)$$

As O'Mahony and Vecchi (2009) show, equation (5) can be used to test for the existence of any conceivable spillover effects arising from dimensions of governance, human capital and physical capital. In a recent study, Bournakis et al. (2017) employ a similar specification to demonstrate the positive spillover effect of human capital and foreign direct investment on output per worker.

MFP growth is obtained as a residual based on growth in output that cannot be explained by growth in inputs. We derive MFP growth from equation (7) below, using Solow's (1957) growth accounting framework and assuming Hicks' neutral technological change, constant returns to scale, decreasing returns to factors of production, optimum utilization of the factors of production, and perfect competition. Solow's growth accounting decomposes output growth into the contributions of factors of production inputs and MFP growth:

$$Y = A\, K^{\alpha}\, L^{1-\alpha} \quad (6)$$

The Cobb–Douglas production function in equation (6) shows that output (Y) depends on the stock of physical capital input (K), labor input (L) and a function of time that allows for neutral technological change (A), which is a shift factor in the production function. We compute real capital stock using the PIM based on Nehru and Dhareshwar (1993), allowing a 5 percent depreciation rate (see Wang & Yao, 2003; World Bank, 2006). Drawing on equation (6), we can express the growth rate of output as

$$\frac{\dot{Y}}{Y} = \alpha \frac{\dot{K}}{K} + (1-\alpha)\frac{\dot{L}}{L} + \frac{\dot{A}}{A} \tag{7}$$

Here, α and $(1-\alpha)$ are the shares of capital and labor, respectively, in total revenue, $\frac{\dot{A}}{A}$ is MFP growth, and the overdot denotes the derivative of the variable with respect to time. Wu (2011) provides a detailed analysis of the two parametric methods used to calculate MFP growth, suggesting that, under Solow's framework, technological progress and MFP growth are treated the same. However, the frontier approach decomposes MFP growth into technical change, efficiency change and scale efficiency.

The results in Table 5 provide some tentative evidence of the positive spillover effects of governance on MFP growth, as the coefficients of the governance indicators are all positive. Again, the estimated coefficients of regulatory quality (REGQU) and government effectiveness (GOVEFF) are significant at conventional levels in the first-difference specification. These findings are consistent with Olson et al. (2000), who show empirically that the quality of governance significantly improves the productivity growth of fast-growing developing countries. As in Table 4, the other four dimensions of governance have the correct positive sign, but are statistically insignificant. One explanation for these results is that the governance data available covers a relatively short period (since 1996). The results suggest that governance reforms need more time to produce a spillover effect. Overall, they show that the evolution of political and economic institutions goes hand in hand with increasing productivity and wealth.

Table 5: Impact of governance on MFP growth, using Barro and Lee measure

	First-difference specification						
	VOACC	**POLST**	**GOVEFF**	**REGQU**	**RLAW**	**CONCORR**	**GOVER**
$\Delta\log(K/L)_t$	0.073	0.074	0.088	0.089	0.063	0.062	0.093
	(0.114)	(0.115)	(0.115)	(0.114)	(0.115)	(0.115)	(0.112)
$\Delta\log(H/L)_t$	0.280**	0.280**	0.285**	0.292**	0.286**	0.288**	0.272**
	(0.127)	(0.127)	(0.128)	(0.130)	(0.127)	(0.129)	(0.124)
$\Delta\log(GOV)_t$	0.009	0.006	0.046**	0.035*	0.005	0.001	0.034
	(0.015)	(0.009)	(0.021)	(0.020)	(0.020)	(0.017)	(0.030)
R-sq.	0.566	0.566	0.578	0.575	0.565	0.565	0.570
Obs.	204	204	204	204	204	204	204
	Levels specification						
	VOACC	**POLST**	**GOVEFF**	**REGQU**	**RLAW**	**CONCORR**	**GOVER**
$\Delta\log(K/L)_t$	0.008	0.001	0.005	0.009	0.007	0.009	0.005
	(0.014)	(0.015)	(0.014)	(0.014)	(0.014)	(0.014)	(0.014)
$\Delta\log(H/L)_t$	0.057	0.081**	0.072*	0.079**	0.068*	0.058	0.068*
	(0.038)	(0.040)	(0.039)	(0.039)	(0.038)	(0.038)	(0.038)
$\Delta\log(GOV)_t$	0.011	0.008*	0.017	-0.03***	0.001	0.014	0.014
	(0.010)	(0.005)	(0.015)	(0.010)	(0.009)	(0.013)	(0.013)
R-sq.	0.514	0.519	0.514	0.524	0.509	0.513	0.513
Obs.	216	216	216	216	216	216	216

Note: *** = significant at 1%, ** = significant at 5%, * = significant at 10%. GOVER is the unweighted governance index derived from the average of the six dimensions of governance. The dependent variable is log ΔMFP (Solow methodology). Heteroskedasticity-consistent standard errors given in parentheses. All equations include country and year dummies.
Source: Authors' calculations.

Table 5 also provides evidence of the positive spillover effects of human capital accumulation in the sampled economies. Our results show that human capital has a positive and significant effect on MFP growth, which is consistent with the positive spillover effect suggested by Bournakis et al. (2017) and Mason et al. (2012). The results of the levels specification are robust to our estimates of the first-difference specification in Table 5. Interestingly, only political stability (POLST) is positive and significant in this specification. Tentatively, these results indicate that political stability is one of the most important factors affecting long-run productivity.

5. Robustness Analysis

To test the robustness of our estimates, we use three different specifications in Table 6, employing primary, secondary and tertiary education in turn as measures of human capital. These findings remain

stable when compared to the results in Table 4. All six dimensions of governance remain positive throughout, although the only ones that are significant are government effectiveness (GOVEFF) and regulatory quality (REGQU). Overall, our results indicate that primary education is positive and significant throughout, emphasizing its importance for developing Asian countries.

Table 6: Impact of governance on ALP growth, first-difference specification

	Primary education as human capital					
	VOACC	**POLST**	**GOVEFF**	**REGQU**	**RLAW**	**CONCORR**
$\Delta\log(K/L)_t$	0.468***	0.469***	0.477***	0.486***	0.457***	0.457***
	(0.102)	(0.104)	(0.105)	(0.101)	(0.105)	(0.104)
$\Delta\log(H/L)_t$	0.222*	0.221*	0.235**	0.230**	0.227*	0.228*
	(0.114)	(0.113)	(0.116)	(0.115)	(0.115)	(0.118)
$\Delta\log(GOV)_t$	0.010	0.006	0.047**	0.035*	0.007	0.002
	(0.015)	(0.009)	(0.022)	(0.020)	(0.020)	(0.017)
R-sq.	0.680	0.680	0.690	0.687	0.680	0.680
Obs.	204	204	204	204	204	204
	Secondary education as human capital					
	VOACC	**VOACC**	**VOACC**	**VOACC**	**VOACC**	**VOACC**
$\Delta\log(K/L)_t$	0.585***	0.577***	0.611***	0.601***	0.570***	0.568***
	(0.115)	(0.112)	(0.114)	(0.114)	(0.112)	(0.109)
$\Delta\log(H/L)_t$	0.077	0.090	0.065	0.086	0.087	0.091
	(0.112)	(0.108)	(0.109)	(0.111)	(0.114)	(0.108)
$\Delta\log(GOV)_t$	0.011	0.007	0.045**	0.0342*	0.006	0.002
	(0.017)	(0.010)	(0.022)	(0.021)	(0.022)	(0.018)
R-sq.	0.673	0.674	0.682	0.680	0.673	0.673
Obs.	204	204	204	204	204	204
	Tertiary education as human capital					
	VOACC	**VOACC**	**VOACC**	**VOACC**	**VOACC**	**VOACC**
$\Delta\log(K/L)_t$	0.579***	0.579***	0.597***	0.598***	0.571***	0.568***
	(0.082)	(0.085)	(0.083)	(0.084)	(0.082)	(0.084)
$\Delta\log(H/L)_t$	0.077	0.076	0.075	0.083*	0.077*	0.079*
	(0.046)	(0.047)	(0.047)	(0.049)	(0.046)	(0.046)
$\Delta\log(GOV)_t$	0.012	0.007	0.045**	0.036*	0.003	-0.003
	(0.016)	(0.009)	(0.021)	(0.020)	(0.021)	(0.017)
R-sq.	0.679	0.679	0.687	0.686	0.678	0.678
Obs.	204	204	204	204	204	204

Note: *** = significant at 1%, ** = significant at 5%, * = significant at 10%. The dependent variable is log ΔALP. Heteroskedasticity-consistent standard errors given in parentheses. All equations include country and year dummies.
Source: Authors' calculations.

Next, we replace Barro and Lee's (2010) measure of human capital with the Cohen–Soto measure: overall, the results remain stable and we conclude once again that government effectiveness and regulatory quality have a positive effect on ALP growth in the sampled Asian economies (see Appendix for results). Similarly, the results show that governance has a positive effect on productivity in these economies. As a final robustness check, we use clustered standard errors via the VCE routine in STATA, which also show that governance has a positive effect on ALP, supporting our earlier findings.[5]

6. Conclusion

Good governance is often regarded as the key to economic growth for developing countries by the World Bank and other international development institutions. In the context of the sampled Asian economies, better governance encourages investment, enhances trade flows, and boosts productivity and economic growth. However, good governance is also considered a 'luxury' that is available only to 'rich' countries (Kaufmann & Kraay, 2002). A key factor explaining the 'East Asian miracle' is the role of dimensions of governance in enhancing productivity and growth. Accordingly, we have attempted to contribute to the debate on the relationship between dimensions of governance, growth and productivity in this region.

We show that government effectiveness and regulatory quality has a positive and significant impact on ALP growth. While the estimated coefficients of the other four governance indicators are generally positive, they are not always significant across different specifications. We argue that this lack of statistical significance may be due to the relatively small time series used, given that institutional development needs time to be effective. When we account for the impact of different levels of education (primary, secondary and tertiary), we find that the contribution of these dimensions of governance remains positive and significant. We also examine the role of these dimensions in enhancing MFP growth. The results support earlier findings (see Hall & Jones, 1999; Acemoglu et al., 2002; Rodrik et al., 2004; Kaufmann et al., 2009) documenting the importance of governance in enhancing economic performance. We also show that human capital has a positive and significant effect on ALP growth. Overall, therefore, we conclude that governance has a positive impact on growth in Asia.

[5] Results available on request.

The Asian economies have achieved strong levels of human capital development over the past 40 years, with average years of schooling among people aged 15 years and above having increased from 2.9 years in 1970 to 7 years in 2010. This progress is tied to the presence of a young population and sturdy improvements in primary and secondary enrolment. However, by 2030, the Asian economies are expected to reach only the 1970 levels of education in advanced economies (Lee & Francisco, 2012). This implies that investing in the quality of governance, together with the quality of education, will be essential to their sustained economic development.

One limitation of this paper is the nonavailability of institutional data at the micro level to explore the mechanism and channels of institution-led development in depth. Future research could examine the role of institutions across regions and gauge whether governance is as important for developing countries as it is for developed ones.

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Appendix

Impact of governance on ALP growth and MFP growth, using Cohen and Soto measure

	First-difference specification: Dependent variable = ALP growth					
	VOACC	**POLST**	**GOVEFF**	**REGQU**	**RLAW**	**CONCORR**
$\Delta \log(K/L)_t$	0.632***	0.614***	0.653***	0.637***	0.609***	0.608***
	(0.151)	(0.137)	(0.142)	(0.138)	(0.136)	(0.134)
$\Delta \log(H/L)_t$	0.016	0.039	0.009	0.037	0.033	0.037
	(0.159)	(0.142)	(0.146)	(0.141)	(0.147)	(0.141)
$\Delta \log(GOV)_t$	0.013	0.008	0.046**	0.035*	0.008	0.002
	(0.019)	(0.010)	(0.023)	(0.021)	(0.022)	(0.017)
R-sq.	0.673	0.673	0.681	0.679	0.672	0.672
Obs.	204	204	204	204	204	204
	First-difference specification: Dependent variable = MFP growth					
	VOACC	**POLST**	**GOVEFF**	**REGQU**	**RLAW**	**CONCORR**
$\Delta \log(K/L)_t$	0.282*	0.264*	0.303**	0.287**	0.259*	0.258*
	(0.151)	(0.137)	(0.142)	(0.138)	(0.136)	(0.134)
$\Delta \log(H/L)_t$	0.016	0.039	0.009	0.037	0.033	0.037
	(0.159)	(0.142)	(0.146)	(0.141)	(0.147)	(0.141)
$\Delta \log(GOV)_t$	0.013	0.008	0.046**	0.035*	0.008	0.002
	(0.019)	(0.010)	(0.022)	(0.021)	(0.022)	(0.017)
R-sq.	0.553	0.554	0.565	0.562	0.552	0.552
Obs.	204	204	204	204	204	204

Note: *** = significant at 1%, ** = significant at 5%, * = significant at 10%. Heteroskedasticity-consistent standard errors given in parentheses. All equations include country and year dummies.
Source: Authors' calculations.