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Abstract

Globalization of financial markets has increased correlation among developed economy markets and made developing economy markets attractive for diversification purposes. Among the developing economy markets are African equity markets, which appear to be the most promising and yet the least studied. Taking the perspective of foreign investors, we estimate the stochastic discount factor (SDF) model for a cross-section of major equity markets in Africa over the period 1997-2009, using the Generalized Method of Moments (GMM). Our findings suggest that real exchange rate risk constitutes a significant time-invariant component of returns in Africa's equity markets. We also find that inflation and nominal exchange rates are separately priced, with time-varying risk premia. Given these findings, international equity investors interested in Africa should hedge their positions against currency risk. Accordingly, African governments should prioritize the development of hedging instruments to increase their equity markets' investability.

JEL classification: G12; G15; F21; F31

Keywords: Currency risk; inflation risk; stochastic discount factor; Africa's equity markets

1. Introduction

Africa's financial markets, many of which have recently instituted liberalization measures, including abolition of capital controls, have attracted increased interest among international investors, largely on account of their low, sometimes negative, correlations with the rest of the world (see e.g., Harvey, 1995), which makes them attractive for portfolio diversification. The importance of Africa's financial markets as international investment destinations is not confined to their low correlations with global markets. It may be explained by many other factors including the fast pace of growth of Africa's economies – in the last decade, six out of ten fastest growing economies globally were in Africa (AfDB, 2013). Many studies have also shown that the performance of Africa's stock markets in the recent past has been remarkable. For instance, Allen et al. (2011) show that stock markets in Africa have recorded impressive risk adjusted returns, with Sharpe indexes as high as 2.47 for Tanzania, 0.49 for Botswana, and 0.42 for Tunisia in 2008; even when converted into US dollar terms, the African region recorded an average annual stock return of 21.8% in 2008.

The potential for diversification benefits presented by Africa's capital markets becomes clearer when one considers the fact that many of the stock markets in the region performed better, on a risk-adjusted basis, than most other markets around the world during the 2007/8 financial crisis with some markets surprisingly generating positive returns (Allen et al., 2011). Similarly, studies have found a positive relationship between risk and expected returns in various stock markets in Africa (Alagidede and Panagiotidis, 2009). This implies that investors who venture into these markets are appropriately rewarded with higher returns for assuming greater risk. Appropriate rewarding of risk appears useful in attracting foreign money to Africa's financial markets. According to AfDB (2013), total external financial flows to Africa reached a historic high of an estimated USD 186.3 billion in 2012, up from USD 158.3 billion in 2011, with the flow of foreign direct investment, portfolio investment, official development assistance and remittances quadrupling since 2001.

In spite of the immense diversification benefits promised by Africa's markets, and financial sector reforms undertaken by many countries in the continent in the last two decades, most of the markets have relatively higher volatility of asset returns driven by thin trading and low liquidity, perceived high political and sovereign risk, relatively weaker regulatory environment and state of development, and lack of high quality accounting data (Abdalla and Murinde, 1997; Appiah-Kusi and Menya, 2003; Alagidede and Panagiotidis, 2009). Further, a set of factors including high risk aversion and home bias among international investors, inadequate information about Africa's markets, global credit tightening, as well as increased volatility of exchange markets (Giovannetti and Velucchi, 2009) characterize many of Africa's financial markets. A combination of these factors is believed to have driven up risk premia in Africa's assets markets, making them unattractive relative to more mature markets. To what extent do these factors actually drive equity risk premia in Africa?

We seek the answer to this question by focusing mainly on one of the most important risk factors – volatility in exchange markets. Our study attempts to establish whether the observed volatility of foreign exchange rates contributes to risk premia in Africa's equity markets. We examine this issue from the perspective of the global investor interested in Africa's equity as a foreign investment. We assume the Purchasing Power Parity principle does not hold in the short-run so that investors resident in different countries have different investment and consumption opportunity sets. In such an investment environment, investors evaluate asset returns in their domestic currency terms and take a keen interest in movements of foreign exchange rates because such movements might affect the domestic value of their foreign (African) equity investment returns. If they perceive exchange rate volatility to be a significant factor affecting their net returns, such investors would price foreign exchange risk, in the sense that it would command a premium that would constitute a part of the overall required rate of return on their African equity investments.

The literature is not very clear on the linkage between foreign exchange rates and stock prices in Africa. Whereas bivariate cointegration tests have not been able to find evidence of the existence of long-run relationship between stock prices and real exchange rates for African countries (Ndako, 2013; Kodongo and Ojah, 2013), a long run relationship exists when the world equity market (the apparent mechanism linking the two financial markets) is introduced (Kodongo and Ojah, 2013). And, although not conspicuous from recent studies (Arief et al., 2009; AfDB, 2011; Aly and Strazicich, 2011), it is believed that economic episodes such as the recent global financial crisis and the European debt crisis adversely affected several African economies through reduced official development assistance, foreign direct investment inflows and worker's remittances to Africa (AfDB, 2013). Since these transmission channels may simultaneously impact the exchange market (supply of foreign exchange) and the stock market (foreign demand for assets), developments in the world financial markets may cause currency risk to be priced in the stock markets.[1] Thus, an understanding of the pricing of currency risk in equity markets is of interest to investors, who would find such knowledge an important ingredient in formulating hedging strategies, and to policy makers, who would utilize such knowledge in formulating and implementing regulatory policies meant to ensure macroeconomic stability and enhance the investability of their financial markets.

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[1] Indeed, AfDB (2013) suggests that a majority of Africa's economies are exposed to developments in Europe, the USA and China, through the exports channel. In 2011, one-third of Africa's merchandise exports went to the European Union (down from 37% in 2006) and more than 11% to the United States against 16% in 2006, while exports to China increased to around 10% of total exports from around 6% in 2006 and exports to India rose to 6% in 2011 from around 4.5% in 2006 (AfDB, 2013: 19)

Yet, Kodongo and Ojah (2011) have found that foreign exchange risk does not command a significant unconditional premium in Africa's major equity markets. This study takes a different approach from that in Kodongo and Ojah (2011). We attempt to determine whether foreign exchange risk is priced in Africa's equity markets in the conditional sense. The conditional pricing approach is motivated by the fact that investors generally form expectations, and hence determine their required rates of return, on the basis of information available at the beginning of their investment decision. Such information, neglected by Kodongo and Ojah (2011), ought to be incorporated by researchers in empirical asset pricing studies (Dumas and Solnik, 1995). We also examine data from a larger sample – ten countries – to provide a broader representation and therefore strengthen generalization. Findings from this study are important to policy in several ways. First, securities markets regulatory bodies and financial markets agencies in African countries should work together towards instituting operationally viable derivative securities markets within their financial jurisdictions, with a view to availing reliable risk management tools to equity markets investors. Second, to minimize volatility in the real exchange rates, and therefore the premium it commands in the stock market, governments should ensure stability in nominal exchange rates and in prices of goods and services in general.

2. Related literature

The pricing of foreign exchange risk in stock markets has been examined by many studies. Early empirical investigations were largely based on the unconditional versions of the models. Conducted in varied markets, some of the unconditional tests fail to detect any evidence of currency risk pricing (Jorion, 1991; Loudon, 1993; Kodongo and Ojah, 2011); others find mixed and inconclusive results (Choi et al., 1998; Iorio and Faff, 2002); yet others present results that are consistent with high degrees of currency risk exposure (Dominguez and Tesar, 2001; Carrieri and Majerbi, 2006). The failure to resolve the issue of currency risk pricing is not desirable from the point of view of investors and corporate finance managers. For instance, company financial managers find it difficult to justify corporate hedging practices in the absence of full knowledge of currency risk pricing. Failure to reject the hypothesis that currency risk is not priced in financial markets implies that investors are not willing to pay a premium for firms with active hedging policies (Dufey and Srinivasulu, 1983; Smith and Stulz, 1985). Inconsistency of research findings by studies based on unconditional asset pricing models has motivated the search for alternative model frameworks. Thus, conditional asset pricing models have recently become popular tools for testing for currency risk pricing.

According to Dumas and Solnik (1995), it is natural to test any asset pricing model in its conditional form. The duo observe that investors' decisions are informed by conditioning information such as interest rates and past equity prices, which empirical tests cannot afford to ignore. Using a parsimonious econometric specification allowing for time-varying currency risk, they conclude that stochastic changes in foreign exchange rates are associated with changes in equity prices and constitute additional sources of risk to investors in the stock markets of Japan, Germany, USA and the UK. De Santis and Gerard (1998) also find currency risk premium to constitute a significant portion of the total risk premium in these four major equity markets. Similar findings have been reported from other financial markets (Doukas et al., 1999; Tai, 1999; MacDonald, 2000; Carrieri, 2001; Kolari et al., 2008).[2] Recently, Aysun and Guldi (2011) compare the performance of linear models and that of non-linear models in explaining currency risk exposure. They find that the proportion of firms with exposures is understated in both emerging markets and the USA under a linear model, but that the frequency of exposure increases when non-linear models are used. Studies that have employed the stochastic discount factor (SDF) model (Zhang, 2006; Cappiello and Panigirtzoglou, 2008) also demonstrate that currency risk premiums are time-varying and account for a significant part of the excess returns on international assets.[3] In an interesting theoretical application of the SDF model, Verdelhan (2010) explains exchange rate risk premia in the context of the uncovered interest rate parity puzzle. The model's features include pro-cyclical real interest rates and time-varying risk aversion.

Some evidence of currency risk pricing is also available from the emerging markets. Phylaktis and Ravazzolo (2004) use a parsimonious multivariate GARCH-M process to find that currency risk premium is time-varying and substantial and forms a big part of the total risk premium; it is bigger and more variable when markets are segmented. Under a partial market segmentation framework, Carrieri et al. (2006) find that local currency risk commands a significant premium, which is, however, smaller, on average, than domestic equity market risk, but which increases substantially during crisis periods, when it can be almost as large as market risk. Using a Value-at-Risk decomposition technique, Sirr et al. (2011) demonstrate significant variation in foreign exchange risk in emerging markets. More recently, Antell and Vaihekoski (2012) use a covariance stationary specification in a multivariate GARCH-M setup to test a conditional international asset pricing model; they find that currency risk is priced in both Swedish and Finnish stock markets, and that the price and the risk premium are lower after the floatation of the currencies, especially for Finland.

The SDF model has also been used successfully to estimate currency risk premia in various countries in the emerging markets. Drobetz et al. (2002) use the SDF model to find that the relative value

[2] Using an alternative methodology, Du and Hu (2012) appear to refute the findings of Kolari et al. (2008).

[3] Time variation in equity risk premia has been attributed to many factors. Jagannathan and Wang (1996) argue that during a recession, financial leverage of troubled firms increases, causing their systematic risk, measured by CAPM beta, to increase. Brooks et al. (1992) also point out that the maturity and growth of firms tend to change the riskiness (beta) of the firm over time. Zhang (2006) observes that the sensitivity to changes in macroeconomic conditions by financially constrained firms may cause their betas to vary over time. Therefore, the market risk premia should be higher at economic downturns to compensate for the adverse risk exposure.

of the US dollar is significantly priced in the emerging markets studied. Poghosyan and Kočenda (2008) find that real factors play a small role in determining foreign exchange risk, while nominal and monetary factors have a significant impact in the post-transition economies of the European Union. Similarly, Poghasyan (2010) find results suggesting that U.S. inflation and consumption growth are important factors driving the risk premium in the Gulf Cooperation Council (GCC) countries.

From the literature it is clear that conditional asset pricing models, if well specified, appear to be better suited than unconditional asset pricing models to succeed in finding currency and inflation risks priced in international equity markets. However, we must point out that the pricing of foreign exchange risk has not been investigated in Africa's equity markets in the conditional sense. Thus, whether the findings from less exotic emerging markets outside Africa can be generalized to Africa's nascent stock markets is still an open empirical question.[4]

We employ the stochastic discount factor (SDF) model with the world equity portfolio and foreign exchange risk as priced factors. Our SDF model allows for time-variation in the pricing kernel and risk premia. We assume that a set of information existing at the beginning of each month conditions investors' decisions. We proxy the conditioning information by four macroeconomic variables believed by international finance researchers to be able to predict equity returns. The conditioning variables include three international variables, namely, the gross returns on the MSCI world equity portfolio, dividend yield on the MSCI world equity portfolio in excess of the Eurodollar deposit rate and the USA term premium. At the local level, we use inflation rates.[5] In addition, consistent with several existing studies, we also use the January dummy. All conditioning variables (except the January dummy) are lagged one period and those that are not stationary in levels are Hodrick and Prescott (1997) filtered. Our investigation covers ten major equity markets in Africa: Botswana, Egypt, Ghana, Kenya, Mauritius, Morocco, Namibia, Nigeria, South Africa and Tunisia, the choice of which we discuss in section 3.3.

Our study adds to the existing literature on conditional currency risk pricing by incorporating a new set of countries. Further, in our robustness check, we estimate the SDF model with real exchange rate decomposed into inflation and nominal exchange rate. Interestingly, we find strong evidence suggesting that both inflation risk and nominal exchange risk are priced with time varying risk premia. This is a material contribution to the literature, which has so far, tended not to study the pricing of inflation in asset markets, especially in the conditional sense. The balance of this paper is structured as follows. Section 3 presents the theoretical model that underpins this study and the empirical procedures, and describes the

[4] Allen et al. (2011) profile Africa's major equity markets in ways that highlight their evolution as well as inform potential international investors about their attractiveness. For additional background information about Africa's national equity markets, see Kenny and Moss (1998) and Nellor (2008).

[5] Other potential variables from Africa's local economies are discussed in the section 3.3.

data. Section 4 presents the primary empirical estimates of the models and explores additional tests as a way of providing robustness checks for the baseline results. Section 5 concludes.

3. Methodology and data

3.1 The theoretical model

The Baseline SDF models

The first order condition for an investor-consumer's portfolio decision problem are:[6]

$$E_t(M_{t+1}R_{i,t+1}) = 1 \tag{1}$$

where $E_t(\cdot)$ is the conditional expectations operator, M_{t+1} is the investor-consumer's intertemporal marginal rate of substitution also known as the pricing kernel, $R_{i,t+1}$ is the gross return anticipated on asset $i, (i = 1, 2, \dots, N)$ at time $t + 1$. Cochrane (2000) demonstrates that equation (1) is equivalent to linear stochastic discount factor models of the form:

$$M_{t+1} = a_t + \boldsymbol{b}_t'\boldsymbol{F}_{t+1} \tag{2}$$

where $\boldsymbol{b}_t$ is a $(k \times 1)$ vector of factor loadings and, $\boldsymbol{F}_{t+1}$ is a $(k \times 1)$ vector of factors. The stochastic discount factor (SDF) model assumes that returns are conditional on the full set of market-wide information, Ω_t, available to investors at time *t*.[7] Since Ω_t is not observable, econometric tests of the SDF model are performed by proxying Ω_t by a set of a few carefully selected and judiciously transformed instrumental variables, $z_t (\in \Omega_t)$, assumed to contain time *t* information. The instrument set can be used to scale the factors, in a situation in which the factors are expected only to conditionally price assets, to give (Cochrane, 1996; Lettau and Ludvigson, 2001; Hodrick and Zhang, 2001):

$$M_{t+1} = a_0 + a_1 z_t + (\boldsymbol{b}_0' + \boldsymbol{b}_1' z_t) \otimes \boldsymbol{f}_{t+1} \tag{3}$$

where $\boldsymbol{f}$ is the vector of scaled and unscaled factors; z_t is a scalar, implying that one conditioning variable is used at a time; $\otimes$ is the Kronecker product (multiply each term in the bracket by every factor). If there are only two factors, we drop the time-varying intercept following Cochrane (1996), and express equation (3) as a five-factor model with parameters b_{11} and b_{12} being allowed to vary over time:

[6] Cochrane (2000) discusses the conditions under which these equations hold. Also see Kodongo (2011) for a more detailed derivation of the models that follow.
[7] Thus, the conditional asset pricing model in equation (1) can also be expressed as $E(M_{t+1}, R_{t+1}|\Omega_t) = 1$.

$$M_{t+1} = a_0 + b_{01}f_{1,t+1} + b_{02}f_{2,t+1} + b_{11}(f_{1,t+1}\, z_t) + b_{12}(f_{2,t+1}\, z_t) \qquad (4)$$

Cochrane (2000) also demonstrates that equation (1) implies a factor model structure of the form:

$$E_t(R_{i,t+1}) = \lambda_0 + \lambda_1\beta_{i,1} + \lambda_2\beta_{i,2} + \ldots + \lambda_k\beta_{i,k} \qquad (5)$$

where the term λ_0 is the price of the zero-beta asset (or the risk-free asset if one exists); λ_j is the risk per unit of the expected return on factor j; $\beta_{i.j}$ $(j = 1, 2, \ldots k)$ is the quantity of risk in asset *i* associated with each factor *j*. By the definition of the covariance, the beta representation in equation (5) is transformed into a relationship of the following form:

$$E(R_{i,t+1}) = \lambda_0 + \lambda_1 \frac{Cov(f_{1,t+1}R_{i,t+1})}{Var(f_{1,t+1})} + \lambda_2 \frac{Cov(f_{2,t+1}R_{i,t+1})}{Var(f_{2,t+1})} + \ldots + \lambda_k \frac{Cov(f_{k,t+1}R_{i,t+1})}{Var(f_{k,t+1})} \qquad (6)$$

3.2 Empirical procedures

Estimation of factor sensitivities

The scaled factor model in equations (3) and (4) can be tested unconditionally by applying the Law of Iterated Expectations[8]. Plugging equation (4) into equation (1), the unconditional sample moment conditions which we evaluate using the Generalized Method of Moments (GMM) are expressed parsimoniously as:

$$g_T(\varphi) = \frac{1}{T}\sum_{t=1}^{T}\{[a_0 + (\boldsymbol{b}_0' + \boldsymbol{b}'z_t) \otimes \boldsymbol{f}_{t+1}]\boldsymbol{R}_{i,t+1} - 1\} = \mathbf{0} \qquad i = 1,2,\ldots,n \qquad (7)$$

where φ is the set of all parameters (a, $\boldsymbol{b}$) to be estimated and $\mathbf{0}$ is an *n*-vector of zeros. Non-zero elements of $\boldsymbol{b}$ indicate the importance of a factor as a determinant of the pricing kernel. We model equation (7) with two factors: the world market risk factor and the currency risk factor. Inclusion of the currency risk factor is informed by Zhang (2006), who finds that exchange risk premiums contribute significantly to the excess returns on international assets, and that the conditional International Capital Asset Pricing Model (ICAPM) with exchange risk performs better than all international asset pricing models that have been hitherto investigated.

[8] The law states that taking an expected value using less information of an expected value that is formed on more information, gives the expected value using less information (Cochrane, 2000). For instance, $E[E_t(X)] = E(X)$.

Estimation of factor risk premia

Using the parameterization in equation (6) and following previous studies (Harvey and Kirby, 1995; Shanken and Zhou, 2007; Iqbal et al., 2010), the sample moment conditions to be evaluated are:

$$g_T(\theta) = \frac{1}{T}\sum_{t=1}^{T}\begin{bmatrix} \boldsymbol{R}_{t+1} - \widehat{\boldsymbol{\mu}}_r \\ \boldsymbol{f}_{t+1} - \widehat{\boldsymbol{\mu}}_f \\ \left(\boldsymbol{f}_{t+1} - \widehat{\boldsymbol{\mu}}_f\right)^2 - \widehat{\boldsymbol{\sigma}}_f^{\boldsymbol{2}} \\ \boldsymbol{R}_{t+1} - \boldsymbol{\iota}\lambda_0 - \sum_{j=1}^{k}\lambda_j \frac{(\boldsymbol{R}_{t+1} - \widehat{\boldsymbol{\mu}}_r)(\boldsymbol{f}_{j,t+1} - \widehat{\boldsymbol{\mu}}_{j,f})}{\widehat{\boldsymbol{\sigma}}_{j,f}^2} \end{bmatrix} = \boldsymbol{0} \tag{8}$$

where a circumflex indicates that the parameter is an estimate; θ is the vector of all parameters; $\boldsymbol{\iota}$ is an *n*-vector of ones; $\boldsymbol{R}_{t+1}$ is an *n*-vector of asset returns with mean vector $\boldsymbol{\mu}_r$; $\boldsymbol{f}_{t+1}$ is a *k*-vector of scaled and unscaled factors with mean vector $\boldsymbol{\mu}_f$; and $\boldsymbol{\sigma}_{j,f}^{\boldsymbol{2}} = Var(\boldsymbol{f}_{j,t+1})$. There are a total of $2(n+k)$ moment conditions in system (8). We estimate the system's parameters through the sequential GMM procedure of Ogaki (1992). Accordingly, we partition the moment conditions $g_T(\theta)$ into two sub-vectors:

$$g_T(\theta) = \frac{1}{T}\sum_{t=1}^{T}\begin{pmatrix} h_{1,t+1}(\theta_1) \\ h_{2,t+1}(\theta_1, \theta_2) \end{pmatrix} \tag{9}$$

The first sub-vector $h_{1,t+1}(\theta_1)$ contains $n+2k$ moment conditions and yields $n+2k$ parameters θ_1 consisting of means of returns and means and variances of factors. The first sub-system is therefore exactly identified so that its GMM estimator $\hat{\theta}_1 = \left(\overline{\boldsymbol{R}}', \overline{\boldsymbol{f}}', \widehat{\boldsymbol{\sigma}}_f^{\boldsymbol{2}\prime}\right)'$ is independent of the weighting matrix. The second sub-vector is defined as $h_{2,t+1}(\theta_1, \theta_2)$ where $\theta_2 = \boldsymbol{\lambda}$ consists of the risk free rate of return and factor risk premia. By plugging $\hat{\theta}_1$ into the last *n* moment conditions and setting $1/T\sum_{t=1}^{T}\left[h_{2,t+1}(\hat{\theta}_1, \theta_2)\right] = 0$, the estimator $\hat{\theta}_2 = \boldsymbol{\lambda}$ is obtained.

Since $\boldsymbol{\lambda}$ is a $k+1(<n)$-vector, the second sub-system is over-identified and the weighting matrix used in the GMM estimation matters. Shanken and Zhou (2007) suggest use of the identity matrix or the inverse of estimates of the variance-covariance matrix of the *n* moment conditions. Harvey and Kirby (1995) demonstrate that the GMM estimator $\hat{\theta}_2$ is fully efficient. Iqbal et al. (2010) point out that the estimator is not subject to errors-in-variables problem because there are no generated regressors employed in the estimation of the risk premia; rather, only the means and variances of returns and factors from the first stage of the sequential GMM process are used. The sequential GMM approach is particularly suitable because, like the GARCH process, it is robust to heteroskedasticity. Further, Hall (1993: 404) explains that under the GMM approach, one need not make an explicit specification of the

data generating process. This can be contrasted with the traditional techniques of estimation, such as maximum likelihood procedures, where a data generating process, typically informed by an explicitly defined probability density function, necessarily forms a part of the estimation process. If the model is misspecified, that is, if the specified distribution function turns out not to accurately reflect the true data generating process, such traditional techniques are likely to give biased parameter estimates: the GMM does not suffer this shortcoming.[9]

Stability tests in conditional asset pricing models

Unconditional asset pricing models have proved ineffective in capturing time-varying risk premiums. This is largely because the theoretical underpinnings of the unconditional models impose strong assumptions on the underlying probability distributions and investors' attitudes toward risk to obtain the time-invariant linear factor structures. Consequently, conditional models have become attractive in empirical asset pricing investigations. However, conditional asset pricing models only work well if they are correctly specified in the sense that the instrumental variables used can correctly capture the dynamics of risk premiums. Ghysels (1998) points out the danger of committing serious pricing errors if the factor risks are inherently misspecified in conditional asset pricing models. He shows that conditional models may have larger pricing errors than unconditional models and attributes this result to structural shifts, the existence of which causes parameter instability.

Garcia and Ghysels (1998) further document the importance of testing for structural changes in the context of emerging markets, especially given the strong political and economic idiosyncrasies that have disrupted these markets in comparison with world markets. To check the structural instability in the SDF parameters, they propose the sup-LM test of Andrews (1993). The null hypothesis for the sup-LM test is that there are no structural shifts so that parameters are stable. This is tested against the alternative that there is a single structural break at some unknown point in time. We compute the sup-LM statistic as the largest of LM statistics computed at 5% increments between 15% and 85% of the sample. The calculated sup-LM statistic is evaluated against critical values in Table 1 of Andrews (2003).

3.3 Data and preliminary analysis

A summary of key facts about the African countries sampled for this study is displayed in Table 1. The countries are sampled on the basis of the "relative trading activity" (a liquidity proxy) of the

[9] For instance, in their work, Moerman and VanDjik (2010), model a GARCH process in which the market risk premium is restricted to be an exponential function of, while the exchange risk premia are a linear function of, the information variables; further, because of the frequent violations of the normality assumption by asset returns, they use a quasi maximum likelihood procedure to estimate the log-likelihood function. Such restrictions need not be imposed, and a particular distribution of returns need not be specified, under the GMM approach.

countries' respective stock markets. Following Allen et al. (2011), we measure relative trading activity as the total value of shares traded on an exchange, as at end of 2009, scaled by the GDP of the exchange's host country for the year 2009. Although liquidity levels for many of the sampled markets are very low by the standards of markets in industrial countries, foreign investor participation in many of the sampled markets is very vibrant. Indeed, for many of the markets, the volume of foreign transactions is, typically, a substantial proportion of total volume within a period: for instance, foreign investor participation at the end of 2011 was about 52% at the Nairobi Securities Exchange (Nyangoro, 2013) and about 56% of the total volume at the Nigerian Stock Exchange.[10]

Table 1 about here

Other than market liquidity, the data in Table 1 show that the countries chosen are also the most trade competitive and have the most viable capital markets in African. Through this set of countries, we also sought to capture a sufficiently representative sample in terms of the level of economic development and spectrum of economic activity. Availability of data and lengths of time over which the data are available also played an important role in informing country choice. Thus, the ten stock markets chosen are some of the most vibrant in Africa, with the longest available stock index data series. The analysis uses aggregate equity market data. All categories of data and all returns are measured in US dollars for the period 1997:1 to 2009:12. This is the period during which foreign investors' participation had been allowed in most of Africa's financial markets (also see official liberalization dates in Table 1) and over which financial time series data are consistently available for the sampled countries. Observations are sampled at monthly intervals.

Three distinct exchange rate regimes can be identified from the ten sampled countries. According to IMF (2008), Morocco, Namibia and Tunisia operate a conventional fixed peg: Namibia has pegged one-for-one against the South African rand while Morocco and Tunisia have pegged against a basket of currencies of key trading partners; Botswana operates a crawling peg (against a basket comprising of the South African rand and the US dollar); South Africa operates and independent float. The rest of the countries *officially* operated a freely floating regime during the study period. However, a country's actual (de facto) exchange rate regime might differ from her officially declared (de jure) exchange rate regime and, in reality, some countries that have declared to have floated their currencies actually only targeted stable exchange rates (Buluba and Otker-Robe, 2002). The duo attribute divergences between stated and actual policies to concerns about political costs of undertaking visible devaluations under a formal peg or

[10] Accessed April 25, 2013 at http://connectnigeria.com/articles/2013/01/21/foreign-investments-in-the-nigerian-stock-exchange-was-56-at-n733-billion-in-2012/

the "fear of floating" to limit the potential impact of large exchange rate exposure on the domestic economy. Indeed, in the new IMF classification that tries to address discrepancies between de jure regimes and de facto regimes, the rest of the sampled countries' exchange rate systems fall under "managed floating with no predetermined path for the exchange rate" (Buluba and Otker-Robe, 2002; IMF, 2008). Thus, other than Namibia, the exchange rates systems for a majority of the countries are flexible enough that risk-averse investors would be concerned about the effect of such flexibility on return from their portfolio investments.

Recent studies (e.g., Patnaik et al., 2010) have, however, demonstrated that Africa, in general, exhibits a shift towards greater exchange rate flexibility. Nonetheless, our model also attempts to address apparent low flexibility in the nominal exchange rates of some of the countries by using real exchange rates. Real exchange rates are preferred because they capture both inflationary forces and nominal exchange rates, both of which, theoretically, have a mutual relationship with cross-border capital flows. Because it incorporates inflation, which responds to domestic production and income levels, among other time-varying macroeconomic fundamentals, real exchange rates, even for countries that operate a conventional fixed peg, are likely to exhibit a great deal of flexibility. From this perspective, we expect that risk-averse international investors would be concerned about the effect of (real exchange rate) fluctuations on returns from their portfolio investments. The relatively more flexible real exchange rate regimes allow us to model and explore the pricing of foreign exchange risk in the sampled equity markets. However, to the extent that real exchange rates may not fully address the issue of flexibility of the exchange rate systems, our tests examine the extent to which foreign exchange restrictions are built into overall return requirements demanded by international equity investors.

The exchange rate is defined as the African currency price of the US dollar so that a positive change indicates depreciation of the African currency. For each African country, the real exchange rate is calculated as the product of the nominal exchange rate with respect to the US dollar and the CPI of the USA relative to the CPI of the foreign country: for instance, the real exchange rate (*RER*) of the Moroccan dirham (MD) with respect to the US dollar ($) is computed as $RER_{MD/\$} = NER_{MD/\$} \times (CPI_{USA}/CPI_{Morocco})$, where *NER* is the nominal exchange rate. To ease comparability with the extant literature, the "Afro" real exchange rate is constructed as in Kodongo and Ojah (2011). Similar exchange rate measures have been used variously in the literature (Vassalou, 2000; Carrieri and Majerbi, 2006). The relative change in the value of foreign currency, s_t, used in parameter estimation, is then calculated as $s_t = \ln S_t - \ ln\, S_{t-1}$, where S_t is the "African" real rate of exchange at time *t*. Inflation is measured as the monthly rate of change in the consumer price index.

Returns on all indexes provided in African currencies are converted into US dollar returns through the nominal exchange rate using the formula: $r_D = \ln(1 + r_f) - \ln(1 + S_D)$, where r_f is the foreign (African) currency-measured return and S_D is the change in the value of the US dollar. As already explained, the study takes the view of a global investor seeking diversification opportunities in foreign (African) equity portfolios. The random rate of return on an index is computed as $\tilde{r}_{i,t} = \ln\ I_{i,t} - \ln\ I_{i,t-1}$, where $I_{i,t}$ is the value of African stock market index, *i*, at time *t*. The gross return on an index is the random rate of return on the index plus one.

Choice of conditioning variables

Dumas and Solnik (1995) explain that instrumental variables can be proxied by endogenous variables, such as financial market variables, that are observed frequently. The general criteria for inclusion of such economic and financial variables are that they must be predictors of return or leading indicators of business cycle (Drobetz et al., 2002). Importantly, conditioning variables should approximate the information set used by investors in setting prices.

Several studies have investigated the predictability of equity returns in the industrial and emerging markets. Fama and French (1988) find that *dividend yields* have systematic forecast power across different time periods and return horizons in the USA. Predictable variation in stock returns is, in turn, tracked by variables commonly used to measure *default* and *term premiums* in bond returns.[11] They conclude that dividend yield and the default spread capture similar variation in expected stock and bond returns. Ang and Bekaert (2007) find that dividend yields predict excess returns only at short horizons and do not have any long-horizon predictive power. Keim and Stambaugh (1986) find, based on the conditional estimates, that there is at best a weak positive *January seasonal* (effect) in the market beta of the difference in returns between small and large firms. Gultekin and Gultekin (1983) find evidence of seasonal patterns of stock returns: the seasonality is manifested in significantly large mean returns in the turn of the tax year, usually January. Systematically higher returns in the month of January, christened *the January effect* in the literature, have also been reported in some emerging market economies (Claessens et al., 1993). Giovannini and Jorion (1987) find that increases in *interest rates* are associated with predictable increases in the volatility of returns in the foreign exchange market and in the US stock market, and that expected returns both in the stock market and in the foreign exchange market are negatively correlated with nominal interest rates. Geske and Roll (1983) find that stock returns are *negatively* related to both expected and unexpected *inflation*. Our preliminary analysis (Table 2) also show that stock returns in the sampled countries are inversely related to local inflation.

[11] The authors define the default-premium as the difference between the yield on a market portfolio of corporate bonds and the yield on Treasury bonds of the same maturity. The term-premium variable is the difference between the Aaa yield (or treasury notes), usually of ten-year maturity, and the one-month Treasury bill rate.

In practice, the selection of instrumental variables can also be guided by previous studies of a similar design. In addition to worldwide factors (such as the dividend yield on world market portfolio and the Eurodollar deposit rate) assumed to apply uniformly across countries, Harvey (1991) considers the following local instruments: lagged own-country equity market index returns, country-specific dividend yield, country-specific short-term interest rates, foreign exchange rate changes, and local maturity spreads. Dumas and Solnik (1995) find that the inclusion of non-dollar interest rates reduces the finite-sample properties of estimates and show that instrument choice is important in determining the success of asset pricing models tests. In contrast, Buckberg (1995) reports substantial improvements in return prediction following inclusion of lagged local market instruments. Accordingly, Buckberg's instruments set includes lagged local dividend yield and the lagged return on the dollar-local currency exchange rate.

Our choice of instrument variables borrows from the existing studies. We make use of ordinary least squares multiple linear regression of gross market index returns on potential predetermined variables frequently used in the literature. The debt markets are not very well developed in the majority of our sampled countries; thus, we could not get enough long term interest rates to develop term premia. Equally, aggregate (market) dividend yields are not available for many of the sampled countries and when available, the series are too short to be of meaningful use. Thus, we investigate a small set of potential local market instruments, including short-term interest rates, lagged local market index returns, exchange rates and local inflation. None of them reports significant coefficients in the predictive regressions. This finding is in contrast with the findings of Garcia and Ghysels (1998) for a range of emerging markets and Iqbal et al. (2010) for Pakistan, perhaps an indication of the uniqueness of Africa's equity markets and why findings from other emerging markets may not be generalized to Africa's markets.

Figure 1 about here

The lagged values of local inflation reports the largest *t*-statistics of all the local variables examined. We therefore include it in the study's instrument variables set. Several worldwide variables report low standard errors in many of the equations, implying good African equity returns predictive power. Thus, in addition to the African country inflation, this study uses lagged values of the following four worldwide instrument variables: the MSCI world equity portfolio returns, the MSCI world equity portfolio dividend yield in excess of the Eurodollar rates, and the USA term premia.[12] Consistent with many asset pricing studies, we also include the January dummy. A key strength of the resulting instrument variables set is that it affords us a direct comparison with the existing literature.

[12] The USA term premium is measured as the difference between monthly returns on ten-year US Treasury bond and monthly returns on three-month US Treasury bills.

The Generalized Method of Moments, used for empirical estimation in this study, rests on the assumption that instrument variables are stationary. Figure 1 display the time series properties of the conditioning variables. The figure shows that the lagged values of the excess MSCI world market dividend yields and USA term premia are not stationary while lagged values of the domestic inflation and gross MSCI world market equity portfolio returns appear stationary. To deal with the problem of non-stationarity and to ensure that the conditioning variables have the ability to capture the time series properties of the risk factors, empirical implementation makes use of their cyclical components of the lagged values of the USA term premia (*UTP*), and excess world dividend yield (*WDX*). The cyclical components of the two conditioning variables are extracted using the Hodrick-Prescott (1997) filter. Visual investigation (Figure 2) shows that the variables, so constructed, are stationary. The Dickey-Fuller statistics for the hypothesis of a unit root, reported in panel B of Table 3, confirm the visual observation.

Figure 2 about here

To demonstrate the ability of the conditioning variables to predict gross returns, we run an ordinary least squares regression of gross returns on equity indexes in the sampled countries on the reconstructed set of conditioning variables. Results are displayed on Table 2. R-squared values indicate that the five conditioning variables explain up to 18.6% of gross returns for Mauritius. The lowest R-squared value is 3.31% for Nigeria. The recorded return predictability is generally consistent with (and in many cases slightly better than) those reported in other equity markets (see e.g., Harvey, 1991; Dumas and Solnik, 1995). The Wald test for the restriction that all coefficients equal zero rejects the null in three countries: Ghana, Nigeria and South Africa. Other than Nigeria, whose stock return predictability is especially low, the excess world equity portfolio dividend yield appears to be an important predictor variable for equity returns in all the countries. Importantly, each of the conditioning variables, excepting African inflation, as already explained, is significant in at least one of the equations.

Table 2 about here

Descriptive statistics

Panel A of Table 3 presents a summary of stock market index returns and their autocorrelations. The assumption of normality appears to be violated by the distributions of returns for all the national market indices as well as the MSCI world equity portfolio index. This is not surprising as evidence suggests that the assumption of normality is frequently violated in asset price returns (see e.g., Brooks et al., 2005). The presence of non-normal return behavior in African capital markets implies that investors are likely to demand more compensation than that implied by return volatility.

Table 3 about here

High positive first order autocorrelation (above 10%), implying high return predictability, is observed for six of the ten countries and the world equity portfolio index. Ghana exhibits the highest persistent return predictability with its autocorrelation function remaining significant up to the ninth lag (0.143). Positive autocorrelation of returns implies that below-average [above-average] returns tend to be followed by other below-average [above-average] returns. Thus, we would expect negative [positive] returns the month following that in which large negative [positive] returns are reported. Consistent with the literature (e.g., Dumas and Solnik, 1995), high and persistent serial dependence is also observed on many of the conditioning variables (panel B). Serial dependence largely dies out by the twenty fourth lag.

Panel C depicts the correlation matrix for the risk factors and conditioning variables. The correlation coefficients are low in all cases, the highest being – 0.234 between the lagged gross MSCI world equity portfolio returns and domestic inflation. Importantly, note that the correlation between nominal exchange rates and inflation, as constructed for this study, is very close to zero. Low correlations among factors and instruments are desired for the efficiency of GMM estimates.

4. Empirical results

We now present empirical results for the estimates of the SDF models discussed in section 3. As already explained, each conditioning variable is incorporated separately into the SDF model. Results are presented separately for each of the different specifications. Inferences are based on the *J*-statistic of Hansen's (1982) optimal iterative GMM.[13] We also provide results for some robustness checks here.

4.1. Parameter estimates

Table 4 contains estimation outputs for parameters of the stochastic discount factor ($\boldsymbol{b}$) as defined in equation (7). These parameters provide information on the importance of each factor in determining the pricing kernel. The table also contains estimates of factor risk premia ($\boldsymbol{\lambda}$) as presented in equation (8). These parameters impart information on the relative importance of each factor in influencing expected returns of equity securities in Africa's capital markets. The model includes two risk factors, namely, the world market equity portfolio returns and real foreign exchange rate changes. Consistent with Cochrane (1996), we use model specifications in which the conditioning variables are not separately included in the

[13] The Hansen and Jagannathan (1997) distance may alternatively be used. However, Hodrick and Zhang (2001) explain that inferences on the validity of a model based on the HJ-distance equal zero are always similar to inferences based on the *J*-statistic.

pricing kernel. Many other authors (Lettau and Ludvigson, 2001; Drobetz et al., 2002; Iqbal et al., 2010), have employed this approach successfully in conditional asset pricing studies in various equity markets.

Table 4 shows that the real exchange risk factor significantly enters the pricing kernel in the equity securities when risk factors are scaled by the lagged MSCI world equity portfolio returns (Panel A) and the January dummy (panel D). In both models, both the constant and the time varying exchange risk factors enter the kernel. However, the time-varying real exchange rate factor enters the pricing kernel in three of the models (panels, A, D and E). In the remaining model specifications, none of the risk factors significantly explains the pricing kernel for equity securities in Africa.

Table 4 about here

Parameter estimates for factor risk premia yield interesting results. The zero-beta asset return lies between 0.5 percent and 1.2 percent per month, which is reasonable for Africa's money markets. If the zero-beta asset can be proxied by Treasury bills, the explanation for the moderate to high zero-beta rate draws from the heavy demand for domestic debt to finance short-term budget deficits in most of the economies studied. The average Treasury bills annualized percentage rate for seven of the countries for which data is available over the study period was 11.45% (approximately, 0.955% per month).[14] Thus, the estimated zero-beta rate is a reasonable estimate of the observed risk-free rates of return. This finding satisfies one of the criteria prescribed by Lewellen et al. (2010), i.e., that conditional asset pricing models must be evaluated against their ability to estimate risk premia that are close to the observed values. We interpret the high *t*-statistics associated with the zero-beta factor to imply that the factor is a suitable proxy for other potential factors affecting equity returns, excluded in our model.

The time-varying component of the world equity market risk factor is significantly priced in four of the five specifications. However, we document mixed findings for the real exchange risk factor. The factor appears to be conditionally priced, with time-varying risk premia, in only two specifications – where risk factors are scaled by lagged returns on the world market equity portfolio (panel A) and United States term premium (panel C). These results do not therefore lead to a clear, unanimous inference about time variation in real exchange risk premia in Africa's stock markets. However, in addition to the two specifications above, the January dummy specification (panel D), also finds the time-invariant real exchange risk factor priced. That makes three out of five specifications, giving us reasonable evidence to infer that the real exchange risk factor generally commands significant *time-invariant* premia in the equity markets studied. This contradicts the unconditional pricing results in Kodongo and Ojah (2011).

[14] The average annualized percentage Treasury bills rates for each country were: Egypt – 8.89; Ghana – 25.01; Kenya – 10.51; Morocco – 4.12; Namibia – 8.806; Nigeria – 12.39 and South Africa – 10.45 (source: International Financial Statistics). No Treasury bills rates were available for the remaining sampled countries for the period.

Diagnostic statistics give a "clean bill of health" to our model specifications. First, the *J*-statistic, the optimal GMM test for over-identifying restrictions, yields *p*-values greater than 10% for all the specifications. Thus, the data do not reject the model specifications at any of the conventional levels of significance. The Sup-LM statistics also indicate that SDF parameter estimates for all the model specification are stable over time. We use the Wald statistic to examine the joint significance of the factor risk estimates: the statistics strongly reject the hypothesis that all factor risk coefficients are zero for all the model specifications tested.

Our findings on foreign exchange risk pricing suggest important implications. First, foreign investors keen to diversify their portfolios in the better performing African equity markets should do so on the understanding that their returns are generally exposed to currency risk and must be hedged against currency fluctuations. However, the pricing of currency risk must be interpreted within context of the particular currency system in each country. Thus, for fixed peg regimes such as, Morocco, Namibia and Tunisia, international equity investors would demand a currency premium to compensate them for unexpected currency devaluations and revaluations in the destination market. The premium may also be justified on the grounds that restrictions imposed on free fluctuations of exchange rates may prevent investors from realizing above average returns from favorable currency movements. For Botswana's crawling peg regimes, patterns of exchange rate movements are still unpredictable as they depend largely on movements in the rand and the dollar, which are in themselves random. However, for international investors whose domestic returns are denominated in the two currencies in the basket, investing in Botswana pula-denominated assets would be less risky and hedging may not be necessary.

For the remaining countries in the sample, exchange rate regimes are sufficiently flexible that their movements are largely random. Many of these countries also frequently suffer inadequate foreign currency reserves further exposing their currencies to irregular fluctuations which may be harmful to investors' unhedged portfolio positions in the short run. Similarly, most of the markets, with the exception of South Africa, have very thin currency trading, relative to more developed markets, which increase the possibility of relatively high currency volatility. As Buluba and Otker-Robe (2002) explain, the depth and size of foreign exchange markets may affect exchange rate volatility. These observations point to a clear need for portfolio hedging against currency fluctuations by international investors.

Figure 3 about here

The foregoing observations also point to the need to broaden financial markets to include more innovative financial products and instruments, such as derivatives, that can aid in the hedging of exposure to currency and other risks encountered by equity investors. Other than South Africa, most of the countries studied here have yet to develop vibrant derivative markets which can aid investors in hedging their exposure against several risk sources. In this regard, African financial markets regulatory authorities should consider the establishment of viable derivative markets within their jurisdictions or expediting the development of these markets if such mechanisms have already been initiated. Although macroeconomic episodes, such as the recent global financial crisis, show that the derivatives tend to encourage speculative activities, which raises volatility in the financial markets with potential destabilization consequences, our recommendation is premised on the desirability of derivatives as tools for risk hedging and price discovery. Thus, African countries that effect this recommendation must also legislate strong regulatory measures to discourage potentially disruptive speculative uses of derivatives.

Next, we evaluate pricing errors of our estimates. Results are displayed in Figure 3. The straight line in each panel is the 45^{o} line, along which all correctly priced assets/portfolios should lie. Looking at the plots, it is clear that pricing errors are fairly large for several countries especially in the specifications in which equity returns are conditioned by lagged values of the United States term premia. The existence of large pricing errors in conditional asset pricing is, however, not novel to this study: it has been reported elsewhere in the emerging equity markets studies (Iqbal et al., 2010) as well as in advanced equity markets studies (Fletcher and Kihanda, 2005; Schrimpf et al., 2007). In general, our pricing error estimates are actually much smaller than those recorded in some less exotic emerging markets (see e.g., Iqbal et al., 2010). Scaling risk factors by the other conditioning variables induces some reduction in pricing errors but still do not fully account for the variation in returns in the African equity markets.

However, Namibia and Botswana are outliers in the better performing (by pricing errors) specifications. For Namibia, leptokurtosis in index returns may explain the large pricing error. Hwang and Satchell (1999) demonstrate that kurtosis is an important factor in modeling emerging market returns. Incidentally, we do not observe a similar error for Nigeria, whose data also exhibit thick tails. Although the descriptive statistics for Botswana do not reveal any serious departures from the "norm," the country has the largest positive average equity returns for the period under investigation. The pricing error is the difference between the average returns and the predicted returns.

In the context of pricing errors, the "best performing" conditioning variable appears to be the African inflation. Thus, of the macroeconomic variables investigated, local inflation seems to be the most capable of predicting future business cycles in the real economies of the African countries studied. This observation appears inconsistent with the earlier OLS regressions in which this variable showed the least ability to predict equity returns.

4.2 Some robustness checks

4.2.1 The pricing of nominal exchange risk and inflation risk

Although empirical investigations of the international asset pricing models with foreign exchange risk abound, as our literature survey shows, relatively little has been done to investigate the role of inflation. Inflation is an important issue for developing countries and is more sensitive to external shocks than other variables such as economic growth (Darne and Ripoll-Bresson 2004). For instance, in recent years, Africa's average inflation rate has increased to around 9% in 2012, from 8.5% in 2011 and 7% in 2010 (AfDB, 2013). The AfDB report documents the policy dilemmas African monetary authorities have to contend with to keep inflationary pressures and fluctuating nominal exchange rates in check. In their study, Iwata and Wu (2006) uses the SDF to find that more than 80% of the volatilities of the currency risk premia can be accounted for by the standard macroeconomic shocks that drive output and inflation. The economic relevance of the influence of expected inflation on stock returns can also be substantial as recently demonstrated by Katzura and Spierdijk (2013).

The key role played by inflation in international asset returns has been recognized by the many theoretical models that suggest a close link between inflation and foreign exchange risk. In their model, Adler and Dumas (1983) incorporate the two variables jointly as the real exchange risk factor. However, this specification, which informed our baseline tests reported in Section 4.1, implies that the prices of inflation and nominal exchange rate risk are restricted to be equal. Relaxing this restriction, leads to a model in which asset returns depend on their sensitivity to both inflation and nominal exchange rate risks (Moerman and van Dijk, 2010). The unconditional pricing of inflation in equity returns has been investigated (Vassalou, 2000; Duarte, 2010). However, our literature survey found only one study (Moerman and van Dijk, 2010) that has investigated inflation as a conditionally priced factor with time varying risk premia in equity returns. The study found inflation risk significantly priced in the G5 (France, Germany, Japan, UK, and USA) equity markets.

Given our findings on the conditional pricing of the real exchange risk, we now attempt to establish whether nominal exchange rate risk and inflation risk are separately priced in the African countries' stock markets. We run the GMM regressions with the same conditioning variables as before. However, the new estimations make use of three risk factors: the world equity market portfolio, the nominal exchange rates and the inflation differential. The inflation differential is computed as:

$$INF = \ln(1 + \varphi_A) - \ln(1 + \varphi_W) \tag{10}$$

where φ_A is the African inflation and φ_W is the world (proxied by the USA) inflation. African inflation is computed as the equal weighted average of the rates of inflation for the ten sampled countries. For each country, inflation is defined as the monthly rate of change in the consumer price index: $\ln(CPI_{t+1}) - \ln(CPI_t)$. We derive the new pricing kernel by extending equation (4) into the following seven-factor constant-weights model. Parameters b_{11}, b_{12} and b_{13} are allowed to vary with time while the remaining parameters capture the time-invariant properties of the risk factors.

$$M_{t+1} = a_0 + b_{01} f_{1,t+1} + b_{02} f_{2,t+1} + b_{03} f_{3,t+1} + b_{11}(f_{1,t+1}\, z_t) + b_{12}(f_{2,t+1}\, z_t) + b_{13}(f_{3,t+1}\, z_t) \quad (11)$$

The sample moment conditions, which we evaluated using the Generalized Method of Moments, are systems (7) and (8), as before. Results, presented in Table 6, show marked improvement in the ability of the conditioning variables to detect the pricing of risk factors. First, all conditioning variables find the time-invariant inflation a significant factor influencing the pricing kernel for equities in the markets studied. Time-varying inflation is significant under two model specifications (panels A and D). Second, the specifications in which the lagged values of the MSCI world equity portfolio returns (panel A), United States term premium (panel C) and domestic inflation (panel E) condition stock returns find the time-varying nominal exchange risk significant for the equity pricing kernel. Further, two specifications (panels A and C) find the time-invariant nominal exchange risk factor significant in the pricing kernel.

Risk premia estimation results show that all the SDF model specifications, but the excess MSCI dividend yield, find inflation risk priced with time-invariant risk premia and all the specifications, excepting lagged African inflation, find inflation risk conditionally priced, with time-varying risk premia. Further, all specifications, with the exception of January dummy, find nominal exchange risk priced, with time-varying risk premia, in Africa's stock markets. Our findings therefore suggest that nominal foreign exchange and inflation risks are separately priced, with time varying risk premia, in Africa's equity markets. The results also show that inflation risk premia are mostly negative. According to Duarte (2010), whose analysis uses the Consumption-CAPM, the negative price of inflation risk arises because high inflation today predicts low growth in future real consumption: periods with positive inflation shocks tend to be bad states of nature and investors are willing to pay insurance in the form of lower mean returns.

Thus, the global investor to Africa's stock markets must be concerned about the impact of these risk sources to the returns on their equity portfolio holdings. Several important policy implications can be gleaned from these findings. First, to reduce instability in inflation rates and hence its impact on equity prices, monetary policy in Africa must be geared towards ensuring stability in the prices of commodities and services. Further, since Verdelhan (2010) has demonstrated that a direct relationship exists between interest rates and foreign exchange risk premia, African governments must put in place policies that would ensure stability in interest rates as a way of controlling the impact of exchange rate fluctuations on equity prices. Similarly, stabilization policies must be put in place to deal with economic shocks that may put short-run pressure on factors that drive exchange rates and inflation.

Table 6 about here

Diagnostic statistics show, first, that the hypothesis that instruments are inappropriate for the models tested is rejected at all conventional levels of significance by the test for over-identifying restrictions (the *J*-statistic); second, the Wald test for the restriction that all coefficient estimates are zero also rejects the null hypothesis in all specifications. Finally, Andrew's (1993) sup-LM statistic rejects the hypotheses of parameter instability in all cases. Thus, the three factor model appears to hold very well.

Figure 4 presents the pricing error plots for the three-factor model specifications. Pricing errors are generally larger than those of the two-factor specifications. The United States term premia still leads the pack of specifications with large pricing errors. Namibia and Botswana are still the outliers for the better performing specifications, the January dummy and African inflation respectively. The African inflation specification is still the "best performer," yielding the lowest pricing errors.

4.2.2 The dynamic performance of the SDF model specifications

Thus far, our discussion has focused on the goodness-of-fit of the SDF model specifications from a cross-sectional perspective. The preceding discussions of pricing errors have also focused largely on risk factor pricing estimates and their associated covariance measure at the neglect of the SDF parameter estimates. The following diagnostic check, proposed by Farnsworth et al. (2002), focuses on the economic magnitudes of pricing errors for particular assets (countries in this paper) from a time series perspective.

Figure 4 about here

The central idea behind the authors' proposal is that if the model works well, then the SDF's pricing errors, $M_{t+1}R_{t+1} - 1$, should not be predictable using any information available at the investors' disposal at time *t*, as proxied by the set of conditioning variables, Z_t. Consistent with equation (4) for the two factor model, and equation (11) for the three-factor model, we define the pricing kernels, M_{t+1}, as

$$M_{t+1} = \hat{b}_0 + \hat{b}_{WLD}R_{WLD} + \hat{b}_{NXR}R_{NXR} + \hat{b}_{WLD\cdot CON}R_{WLD\cdot CON} + \hat{b}_{RXR\cdot CON}R_{RXR\cdot CON} \tag{12}$$

$$M_{t+1} = \hat{b}_0 + \hat{b}_{WLD}R_{WLD} + \hat{b}_{CXR}R_{CXR} + \hat{b}_{INF}R_{INF} + \hat{b}_{WLD\cdot CON}R_{WLD\cdot CON} + \hat{b}_{CXR\cdot CON}R_{CXR\cdot CON} + \hat{b}_{INF\cdot CON}R_{INF\cdot CON} \tag{13}$$

where *WLD* represents the world market equity portfolio, *RXR* denotes the change in real exchange rate *CXR* is the change in the nominal exchange rate, *INF* is the inflation differential and *CON* is a conditioning variable; R_k is the gross return on factor k, or interaction between factor and conditioning variable, k, at time $t+1$; a circumflex above a coefficient denotes the GMM estimate of the coefficient.

Table 5 about here

Farnsworth et al. (2002) demonstrate that the standard deviations of the fitted values of a regression of the model pricing errors, $M_{t+1}R_{t+1} - 1$, on instrumental variables set, Z_t, should explain how well a particular model specification accounts for predictable variation in returns.[15] The smaller the sample standard deviation, the better is the model specification's ability to "explain" return variation. The one strength of this diagnostic check is that standard deviation measure does not place any penalty on model specifications that get the average return wrong.

We perform this test for each of the ten countries in the sample; the results are presented in Table 5. In terms of the average standard deviation, the table shows that the specification with *LIN* as the conditioning variable appears to perform the best in capturing the time series predictability of gross returns on national equity indexes for the two-factor (real exchange risk) model. Thus, consistent with inference drawn from the *J*-statistic, the African inflation still proves to be the "best performer" – i.e., it is the most able, of the instrumental variables, to capture dynamics of the business cycles in the economies examined. It is worth noting that one of the specifications with better cross-sectional performance, *JAN*, also turns out to be the "worst performer" from a time-series perspective. For the three-factor model (nominal exchange rates and inflation), the United States tem premium (*UTP*) is the best performer from a time series perspective. This confirms Harvey's (1991) contention that USA macroeconomic variables are good predictors of equity returns in other countries. Incidentally, UTP was also the worst performer on the cross-sectional perspective. The latter observation is demonstrates that the Farnsworth et al. (2002) methodology is not biased by the ability of a model specification to correctly capture average returns.

5. Conclusions

Africa's financial markets have recently attracted increased interest among international investors, largely on account of their low, sometimes negative, correlations with the rest of the world, which makes them attractive for portfolio diversification and also perhaps due to the fact that Africa has some of the fastest growing economies in the world today. Studies have also found a positive relationship between risk and expected returns in various stock markets in Africa. Total external financial flows to Africa

[15] Schrimpf et al. (2007) and Iqbal et al. (2010) implement a variant of this approach in which the vector of excess asset returns, rather than gross returns, is employed in the regressions. Consistent with Farnsworth et al. (2002), they both find that conditional asset pricing models struggle to capture time-varying predictability of stock returns. Indeed, differences in results are not expected provided that the same risk-free rate of return is used to compute excess returns for all portfolios/assets.

reached a historic high of an estimated USD 186.3 billion in 2012, up from USD 158.3 billion in 2011, with the flow of foreign direct investment, portfolio investment, official development assistance and remittances quadrupling since 2001 (AfDB, 2013). Yet most of Africa's financial markets have relatively higher volatility of asset returns driven by thin trading and low liquidity, perceived high political and sovereign risk, relatively weaker regulatory environment and state of development, and lack of high quality accounting data. Further, inadequate information about Africa's markets, global credit tightening, as well as increased volatility of foreign exchange markets, characterize many of Africa's financial markets. A combination of these factors is believed to have driven up risk premia in Africa's assets markets, making them unattractive relative to more mature markets.

To what extent do these factors actually drive equity risk premia in Africa? Our study uses the SDF model in an attempt to establish whether the observed volatility of foreign exchange rates contributes to risk premia in Africa's equity markets. We use a two-factor model with the world market equity portfolio and (real) foreign exchange rate changes as risk factors. Different specifications of the two-factor model, in which the factors are scaled by conditioning variables, are examined. Scaling factors helps to capture time variation in risk premia. To approximate the time-varying risk premia, five conditioning variables, namely, the gross return on the MSCI world equity portfolio, MSCI world equity portfolio dividend yield in excess of the Eurodollar rate, the USA term premium, the January dummy and African inflation are used.

As a robustness check, we run a three factor model with nominal exchange rates and inflation differential separately as priced factors. None of the specifications is rejected but pricing errors are larger. We also run a test of the dynamic performance of the model specifications following Farnsworth et al. (2002). The test finds that the specification that prices average returns best (African inflation) also performs the best in capturing the time series predictability in asset returns for the two-factor model. However, the United States term premium performs better than other conditioning variables for the three-factor model. These findings give contradictory signals – while the two-factor model concurs with Garcia and Ghysels (1998) that local variables (in our case, inflation) perform better in predicting equity returns in local markets, the three-factor model results appear to conform to the Harvey's (1991) argument that USA macroeconomic variables (USA term premia) can help explain equity returns in other countries.

In the main, our empirical results suggest that the world market equity factor does not, in general, significantly contribute to the pricing kernel in Africa's equity markets. However, the world market factor is found to have significant time-varying risk premia. Further, there is evidence suggesting that foreign exchange rate changes have significant time-invariant risk premia under the two factor model and both are priced conditionally and unconditionally under the three-factor structures. Finally, we also find strong evidence to suggest that inflation is priced both conditionally and unconditionally in Africa's equity markets. These findings suggest some implications. First, international investors in Africa's equity markets must ensure that their positions are hedged against currency risk. Second, security markets regulators in African countries should establish well-functioning derivative markets to help investors hedge their positions against adverse shocks emanating from factors exogenous to the stock markets. Finally, African governments should put measures in place to ensure price stability. These measures, collective, will enhance the investability of Africa's financial markets.

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Table 1: Some facts about the sampled countries

Country	Market capitalization (billions of US dollars)[1]	Relative trading activity (%)[2]	Restriction of capital flows[3]	Official Market Liberalization year[4]	Country Competitiveness index[5]	Foreign participation[6]
Botswana	59.1	0.89	5.7	N/A	4.2	Yes
Egypt	90.1	28.04	4.4	1992	4.0	Yes
Ghana	10.9	0.37	4.2	1990	3.6	Yes
Kenya	11.8	1.65	4.4	1995	3.8	Yes
Mauritius	4.8	3.83	6.2	1994	4.2	Yes
Morocco	63.0	32.38	3.6	1988	4.1	Yes
Namibia	0.8	0.24	3.9	N/A	4.0	Yes
Nigeria	47.0	2.71	4.1	1995	3.8	Yes
S. Africa	790.5	119.76	3.7	1996	4.4	Yes
Tunisia	9.1	3.18	4.4	1995	4.6	Yes

[1] From the websites of the relevant stock exchanges, as of end of December 2009.
[2] Total value of shares traded at an exchange, in 2009, as a proportion of the country's 2009 GDP, reported in Allen et al. (2011). Of the other African countries with these data, Cote d'Ivoire scored 0.58. However, available data series for Cote d'Ivoire is not long enough that we could include it in our sample.
[3] As of 2009, on the scale 1 (highly restricted by law) to 7 (not restricted by law) (World Economic Forum, 2009a).
[4] Dates are those reported in Bekaert et al. (2003) except for Ghana and Mauritius whose liberalization dates are from, respectively, Adam (2009) and the African Development Bank website: http://www.afdb.org/fileadmin/uploads/afdb/Documents/Publications/24108404-EN-MAURITIUS.PDF
[5] As of 2009, on a scale 1 (least competitive) to 7 (most competitive). (World Economic Forum, 2009a).
[6] See also Allen et al. (2011). For Botswana, Namibia and Mauritius, there are restrictions on the amount of foreign ownership for some "strategic" companies.

Table 2 Regression of gross returns on lagged conditioning variables

	Intercept	WLR	WDX	UTP	INF	JAN	Wald	R^2
Botswana	0.643 (6.88)	0.361*** (3.99)	-7.935* (-1.74)	9.101 (1.11)	-1.499 (-1.38)	-0.030* (-1.87)	4.52 [0.0004]	0.1266
Egypt	0.682 (4.66)	0.380*** (2.68)	17.047** (2.39)	-19.577 (-1.53)	-1.26 (-0.74)	0.030 (1.19)	3.41 [0.0046]	0.0985
Ghana	1.045 (10.40)	-0.045 (-0.46)	5.444 (1.11)	2.533 (0.29)	-0.276 (-0.24)	0.010 (0.56)	1.62 [0.1528]	0.0492
Kenya	0.885 (7.10)	0.182 (1.51)	17.414*** (2.87)	-25.539** (2.34)	-2.311 (1.60)	-0.015 (-0.69)	2.59 [0.0243]	0.0766
Mauritius	0.615 (7.45)	0.415*** (5.19)	8.948** (2.22)	-13.043* (-1.80)	0.444 (0.46)	0.010 (0.75)	7.13 [0.0000]	0.1861
Morocco	0.899 (9.03)	0.137 (1.42)	0.095** (2.08)	-22.458** (-2.57)	1.612 (1.39)	0.012 (0.73)	2.51 [0.0285]	0.0744
Namibia	0.623 (3.65)	0.425** (2.58)	14.693* (1.77)	-18.765 (-1.26)	0.332 (0.17)	-0.021 (-0.75)	2.24 [0.0481]	0.0670
Nigeria	0.897 (3.93)	0.114 (0.52)	7.360 (0.66)	-9.788 (0.49)	2.431 (0.91)	-0.067* (-1.74)	1.07 [0.3759]	0.0331
South Africa	0.883 (5.91)	0.169 (1.17)	15.948** (2.19)	-20.113 (-1.54)	0.104 (0.06)	-0.009 (-0.36)	1.45 [0.2046]	0.0443
Tunisia	1.034 (8.88)	0.016 (0.15)	19.116*** (3.37)	-25.718** (-2.52)	0.522 (0.39)	-0.001 (-0.04)	2.55 [0.0264]	0.0755

The table uses data for the period 1997:1 through 2009:12. The two values reported in the body of the table are, respectively, the coefficient of the explanatory variable and its corresponding *t*-statistic (in parentheses). The dependent variables are the respective gross returns on the local country equity market index, measured in US dollars. The explanatory variables are the January dummy (*JAN*) and lagged values of returns on the world market portfolio (*WLR*), monthly dividend yield on the world market portfolio in excess of the Eurodollar rate (*WDX*), the monthly USA term premium (*UTP*), the one-month African inflation (*LIN*). "Wald" is the F-statistic from the test that all regression coefficients are zero; its accompanying *p*-value is in square brackets. R^2 is the coefficient of determination.*, **, and *** denote statistical significance at 10%, 5% and 1% respectively.

Table 3 Summary statistics for gross returns on national market indexes, and conditioning variables

	Mean	Std dev.	Skew	Kurt	Jarque -Bera	Autocorrelations ρ_1	ρ_2	ρ_4	ρ_6	ρ_{12}	ρ_{24}
Panel A: Gross returns on national market indices											
Botswana	1.0157	0.0578	-0.077	0.741	33.33	0.36[a]	0.16[a]	0.08	0.08	-0.05	-0.04
Egypt	1.0101	0.0892	-0.007	-0.287	70.24	0.38[a]	0.21[a]	0.16[a]	0.04	0.07	0.08
Ghana	0.9945	0.0595	-0.524	3.516	8.88	0.37[a]	0.27[a]	0.21[a]	0.17[a]	-0.04	-0.05
Kenya	0.9959	0.0749	-0.936	6.360	96.15	0.21[a]	0.02	0.17	0.04	0.06	0.03
Mauritius	1.0063	0.0529	-0.869	8.623	>99.99	0.42[a]	0.29[a]	0.19[a]	-0.01	-0.16	-0.07
Morocco	1.0062	0.0598	0.048	1.200	21.11	0.08	0.12	0.05	0.11	0.14	0.05
Namibia	1.0031	0.1018	-1.828	10.501	>99.99	0.18[a]	0.01	0.01	-0.02	-0.02	-0.08
Nigeria	1.0000	0.1343	-7.156	76.044	>99.99	0.07	0.02	0.07	0.06	-0.03	-0.06
South Africa	1.0053	0.0883	-1.092	2.336	33.90	0.07	-0.04	-0.04	0.02	0.03	-0.13
Tunisia	0.9967	0.0700	-0.033	8.418	>99.99	0.08	-0.02	-0.03	0.08	0.13	0.07
MSCI World	1.0023	0.0482	-1.074	2.327	32.92	0.22	0.00	0.14	-0.06	0.07	0.01
Panel B: Instrument variables											
			ADF								
WLR			-9.928		59.8	0.21[a]	0.00	0.14	-0.06	0.07	0.00
WDX			-8.509		>99.9	0.51[a]	0.09	-0.38[a]	-0.28[a]	0.20	0.04
UTP			-9.194		2.25	0.52[a]	-0.02	-0.23[a]	-0.14	-0.14	0.14
LIN			-7.872		16.2	0.42[a]	0.22[a]	-0.05	-0.18[a]	0.29[a]	0.27[a]

Panel C: Correlations

	Conditioning variables				Risk factors		
	LIN	UTP	WDX	WLR	INF	CNR	CRR
WLD	0.060	-0.030	0.301	0.216	-0.024	-0.179	-0.178
CRR	-0.030	-0.028	-0.132	0.026	-0.083	0.980	
CNR	-0.053	-0.024	-0.170	0.003	0.093		
INF	-0.014	0.083	-0.205	-0.234			
WLR	0.111	-0.045	0.080				
WDX	0.233	0.106					
UTP	-0.008						

[a] indicates that the autocorrelation is significantly different from zero. "Skew" and "Kurt," respectively, stand for skewness and kurtosis. *WLR, WDX, UTP*, and *LIN* represent, respectively, the lagged values of gross returns on the world equity portfolio, the Hodrick-Prescott (1987) filtered lagged values of the MSCI world equity portfolio dividend yield in excess of the Eurodollar rates, the USA term premia. Gross returns on the world market equity portfolio (WLD), change in the nominal exchange rates (CNR), change in the real exchange rates (CRR) and inflation differential (INF) respectively are the risk factors. The inflation differential is used later together with nominal exchange rates to check the model's robustness. All returns are in US dollars; they cover the period 1997:1 to 2009:12. ADF is the augmented Dickey-Fuller test statistic for the hypothesis of a unit root. Critical values of the ADF statistic are -3.472, -2.880, and -2.577 at 1%, 5% and 10% levels respectively.

Table 4 Parameter estimates for the SDF model specifications

Panel A: Factors scaled by returns on MSCI world equity portfolio (WLR)					
SDF Parameters	Constant	b_{WLD}	b_{RXR}	$b_{WLD \cdot WLR}$	$b_{RXR \cdot WLR}$
Coefficient	5.236*	-0.370	-0.469***	-0.553	0.444***
(t-value)	(1.89)	(-1.03)	(-2.96)	(-0.51)	(2.94)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	$\lambda_{WLD \cdot WLR}$	$\lambda_{NXR \cdot WLR}$
Coefficient	1.009***	-0.0005	0.336***	0.029*	-3.404***
(t-value)	(374.57)	(-0.05)	(3.08)	(1.74)	(-3.06)
Model Tests			J-Statistic	Wald	Sup-LM
Chi-square			17.61	52.33	5.78
(p-value)			[0.2839]	[0.0001]	

Panel B: Factors scaled by world equity dividend yield in excess of Eurodollar rate (WDX)					
SDF Parameters	Constant	b_{WLD}	b_{RXR}	$b_{WLD \cdot WDX}$	$b_{RXR \cdot WDX}$
Coefficient	-0.066	1.059	-0.021	-1.14	0.039
(t-value)	(-0.03)	(0.49)	(-0.93)	(-0.85)	(0.67)
Factor Risk Premia	Constant	λ_{WLD}	λ_{RXR}	$\lambda_{WLD \cdot WDX}$	$\lambda_{RXR \cdot WDX}$
Coefficient	1.009***	-0.006	0.156	0.027	-0.006
(t-value)	(366.22)	(-1.35)	(1.11)	(1.47)	(-0.08)
Model Tests			J-Statistic	Wald	Sup-LM
Chi-square			19.73	19.57	7.58
(p-value)			[0.1829]	[0.0006]	

Panel C: Factors scaled by USA term premium (UTP)					
SDF Parameters	Constant	b_{WLD}	b_{RXR}	$b_{WLD \cdot UTP}$	$b_{RXR \cdot UTP}$
Coefficient	0.684	0.315	-1.020	-0.338	-0.268
(t-value)	(0.25)	{0.12)	(-1.48)	(-0.82)	(-1.10)
Factor Risk Premia	Constant	λ_{WLD}	λ_{RXR}	$\lambda_{WLD \cdot UTP}$	$\lambda_{RXR \cdot UTP}$
Coefficient	1.010***	0.016**	2.448***	0.211***	0.747***
(t-value)	(251.24)	(2.10)	(3.97)	(3.23)	(3.70)
Model Tests			J-Statistic	Wald	Sup-LM
Chi-square			15.37	21.09	3.93
(p-value)			[0.4252]	[0.0003]	

Panel D: Factors scaled by January Dummy (JAN)					
SDF Parameters	Constant	b_{WLD}	b_{RXR}	$b_{WLD \cdot JAN}$	$b_{RXR \cdot JAN}$
Coefficient	10.566*	-0.962*	-0.525*	0.260**	0.049*
(t-value)	(1.95)	(-1.77)	(-1.85)	(2.11)	(1.71)
Factor Risk Premia	Constant	λ_{WLD}	λ_{RXR}	$\lambda_{WLD \cdot JAN}$	$\lambda_{RXR \cdot JAN}$
Coefficient	1.012***	0.039***	0.069***	0.127***	-0.020
(t-value)	(345.59)	(5.92)	(4.37)	(6.11)	(-1.28)
Model Tests			J-Statistic	Wald	Sup-LM
Chi-square			10.08	46.93	13.56
(p-value)			[0.8144]	[0.0001]	

Panel E: Factors scaled by lagged African inflation (LIN)					
SDF Parameters	Constant	b_{WLD}	b_{RXR}	$b_{WLD \cdot LIN}$	$b_{RXR \cdot LIN}$
Coefficient	1.456	-0.0430	0.208	-0.015	0.046*
(t-value)	(0.42)	(-0.12)	(1.49)	(-0.20)	(1.66)
Factor Risk Premia	Constant	λ_{WLD}	λ_{RXR}	$\lambda_{WLD \cdot LIN}$	$\lambda_{RXR \cdot LIN}$
Coefficient	1.005***	0.011*	0.002	0.014***	-0.024
(t-value)	(399.46)	(1.77)	(1.52)	(4.32)	(-0.40)
Model Tests			J-Statistic	Wald	Sup-LM
Chi-square			21.32	33.15	12.50
(p-value)			[0.1270]	[0.0001]	

The table uses monthly nominal gross returns for the period from 1997:1 to 2009:12. Returns are denominated in US dollars. The table reports GMM estimates of parameters of the stochastic discount factor model (b) and factor risk premia (λ). The reported t-statistics are robust to heteroskedasticity and autocorrelation. J-statistic is Hansen's (1982) test of overidentifying restrictions; Wald1 is the joint Wald test that all factor pricing parameters equal zero; Wald2 is the joint Wald test that parameters λ_{NXR} and $\lambda_{NXR.CON}$ equal zero (where *CON* is a specific conditioning variable). *WLD* and *RXR* stand for gross monthly returns on MSCI world equity portfolio and monthly real foreign exchange rates respectively. The foreign exchange rate factor is computed as in Kodongo and Ojah (2011). All conditioning variables, except the January dummy, are lagged one period. Numbers in square brackets are *prob*-values of the test statistics. Sup-LM is Andrews (1993) test of structural stability of the parameters of the SDF model. *, ** and *** indicate statistical significance at 10%, 5% and 1% respectively. Critical values of sup-LM statistics obtained from Table I of Andrews (2003) are: 1% - 20.47; 5% - 16.87 and 10% - 17.78.

Table The dynamic Performance of the SDF model specifications

	Average standard deviation	
Risk factors scaled by	Two-factor model	Three-factor model
WLR	0.0292	0.3764
WDX	0.0227	0.0513
UTP	0.0162	0.0030
JAN	0.0799	0.7167
LIN	0.0077	0.0141

The table uses monthly nominal gross returns, denominated in US dollars, for the period from 1997:1 to 2009:12. The table reports results of the tests for dynamic model specification performance following Farnsworth et al. (2002). The model pricing errors, $M_{t+1}(\hat{b})R_{t+1} - 1$, for each model specification are regressed against the relevant conditioning variable. The standard deviations of the resulting fitted values are then obtained in each case. The conditioning variables used are the gross returns on the MSCI world equity portfolio (WLR), dividend yields on the MSCI world equity portfolio in excess of the Eurodollar deposit rates (WDX), the USA term premium (UTP), the January dummy (JAN) and African inflation rates (LIN). All conditioning variables, except the January dummy are lagged one period.

Table 6 Parameter estimates with nominal exchange rates and inflation as separate risk factors

Panel A: Factors scaled by returns on MSCI world equity portfolio (WLR)							
SDF Parameters	Constant	b_{WLD}	b_{NXR}	b_{INF}	$b_{WLD \cdot WLR}$	$b_{NXR \cdot WLR}$	$b_{INF \cdot WLR}$
Coefficient	7.289**	-0.636	-2.052**	-19.161***	-6.770***	1.899**	2.224***
(t-value)	(2.10)	(-0.44)	(-2.04)	(-3.14)	(-2.76)	(1.98)	(3.19)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	λ_{INF}	$\lambda_{WLD \cdot WLR}$	$\lambda_{NXR \cdot WLR}$	$\lambda_{INF \cdot WLR}$
Coefficient	1.004***	0.001	0.129***	0.153***	0.019	-0.129***	-1.817***
(t-value)	(351.58)	(0.19)	(2.73)	(5.23)	(1.50)	(-2.68)	(-5.73)
Model Tests		J-Statistic		Wald		Sup-LM	
Chi-square		9.22		98.79		19.40	
(p-value)		[0.7563]		[<0.0001]			
Panel B: Factors scaled by world dividend yield in excess of Eurodollar deposit rate (WDX)							
SDF Parameters	Constant	b_{WLD}	b_{NXR}	b_{INF}	$b_{WLD \cdot WDX}$	$b_{NXR \cdot WDX}$	$b_{INF \cdot WDX}$
Coefficient	-1.601	2.924*	0.000	-0.765**	-0.085**	-0.142	-0.237
(t-value)	(-0.94)	(1.71)	(0.01)	(-2.52)	(-0.17)	(-0.21)	(-0.76)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	λ_{INF}	$\lambda_{WLD \cdot WDX}$	$\lambda_{NXR \cdot WDX}$	$\lambda_{INF \cdot WDX}$
Coefficient	1.011***	-0.001	0.144	-0.064	0.104***	-0.009*	-0.065***
(t-value)	(348.27)	(-0.17)	(0.90)	(-0.99)	(3.50)	(-1.06)	(-3.17)
Model Tests		J-Statistic		Wald		Sup-LM	
Chi-square		14.18		55.71		10.15	
(p-value)		[0.3614]		[<0.0001]			
Panel C: Factors scaled by USA term premium (UTP)							
SDF Parameters	Constant	b_{WLD}	b_{NXR}	b_{INF}	$b_{WLD \cdot UTP}$	$b_{NXR \cdot UTP}$	$b_{INF \cdot UTP}$
Coefficient	0.535	0.165	-0.120**	0.937*	0.492	-0.597*	-0.689
(t-value)	(0.22)	(0.07)	(-2.15)	(1.71)	(1.08)	(-1.87)	(-0.79)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	λ_{INF}	$\lambda_{WLD \cdot UTP}$	$\lambda_{NXR \cdot UTP}$	$\lambda_{INF \cdot UTP}$
Coefficient	1.003***	0.019**	1.676**	-0.379***	-0.024***	0.488*	0.179***
(t-value)	(226.36)	(3.02)	(2.11)	(-5.29)	(-0.30)	(1.84)	(3.09)
Model Tests		J-Statistic		Wald		Sup-LM	
Chi-square		10.60		41.94		9.22	
(p-value)		[0.6447]		[<0.0001]			
Panel D: Factors scaled by January dummy (JAN)							
SDF Parameters	Constant	b_{WLD}	b_{NXR}	b_{INF}	$b_{WLD \cdot JAN}$	$b_{NXR \cdot JAN}$	$b_{INF \cdot JAN}$
Coefficient	2.167	-1.986	-0.023	2.088**	2.618***	-0.328	-3.875***
(t-value)	(0.49)	(-0.45)	(-0.08)	(2.33)	(5.14)	(-0.12)	(2.40)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	λ_{INF}	$\lambda_{WLD \cdot JAN}$	$\lambda_{NXR \cdot JAN}$	$\lambda_{INF \cdot JAN}$
Coefficient	1.005***	0.006	1.468	-0.179**	0.132***	-0.013	0.055
(t-value)	(332.99)	(3.61)	(1.12)	(-2.04)	(5.28)	(-1.14)	(1.05)
Model Tests		J-Statistic		Wald		Sup-LM	
Chi-square		11.97		166.73		6.86	
(p-value)		[0.5303]		[<0.0001]			

Panel E: Factors scaled by lagged African inflation (LIN)							
SDF Parameters	Constant	b_{WLD}	b_{NXR}	b_{INF}	$b_{WLD \cdot LLM}$	$b_{NXR \cdot LLM}$	$b_{INF \cdot LLM}$
Coefficient	4.483	-4.548	0.025	2.807***	-0.909*	8.824***	0.984
(t-value)	(1.52)	(-1.53)	(0.77)	(3.19)	(-1.52)	(2.77)	(0.93)
Factor Risk Premia	Constant	λ_{WLD}	λ_{NXR}	λ_{INF}	$\lambda_{WLD \cdot LLM}$	$\lambda_{NXR \cdot LLM}$	$\lambda_{INF \cdot LLM}$
Coefficient	1.001***	0.005	0.136	-0.335***	0.002	-0.021***	-0.018
(t-value)	(315.78)	(1.09)	(0.64)	(-4.61)	(0.78)	(-4.12)	(-0.42)
Model Tests		J-Statistic		Wald		Sup-LM	
Chi-square		11.76		45.26		10.94	
(p-value)		[0.5472]		[<0.0001]			

The table uses monthly nominal gross returns for the period from 1997:1 to 2009:12. Returns are denominated in US dollars. The table reports GMM estimates of parameters of the stochastic discount factor model (b) and factor risk premia (λ). The reported t-statistics are robust to heteroskedasticity and autocorrelation. J-statistic is Hansen's (1982) test of overidentifying restrictions; Wald is the joint Wald test that all factor pricing parameters equal zero. *WLD*, *NXR* and *INF* stand for gross monthly returns on MSCI world equity portfolio, monthly nominal foreign exchange rates and inflation differentials respectively. The foreign exchange rate factor is computed as in Kodongo and Ojah (2011). Inflation differential is the log difference of African inflation and USA (world proxy) inflation. Inflation for each country is the first difference of the country's CPI. African inflation is the equal weighted average of inflation of African countries. All conditioning variables, except the January dummy, are lagged one period. Numbers in square brackets are *prob*-values of the test statistics. Sup-LM is Andrew's (1993) test for parameter stability. Critical values of sup-LM statistics obtained from Table I of Andrews (2003) are: 1% - 26.23; 5% - 21.84 and 10% - 19.69. *, ** and *** indicate statistical significance at 10%, 5% and 1% respectively.

Figure 1 Time series behavior of instrument variables

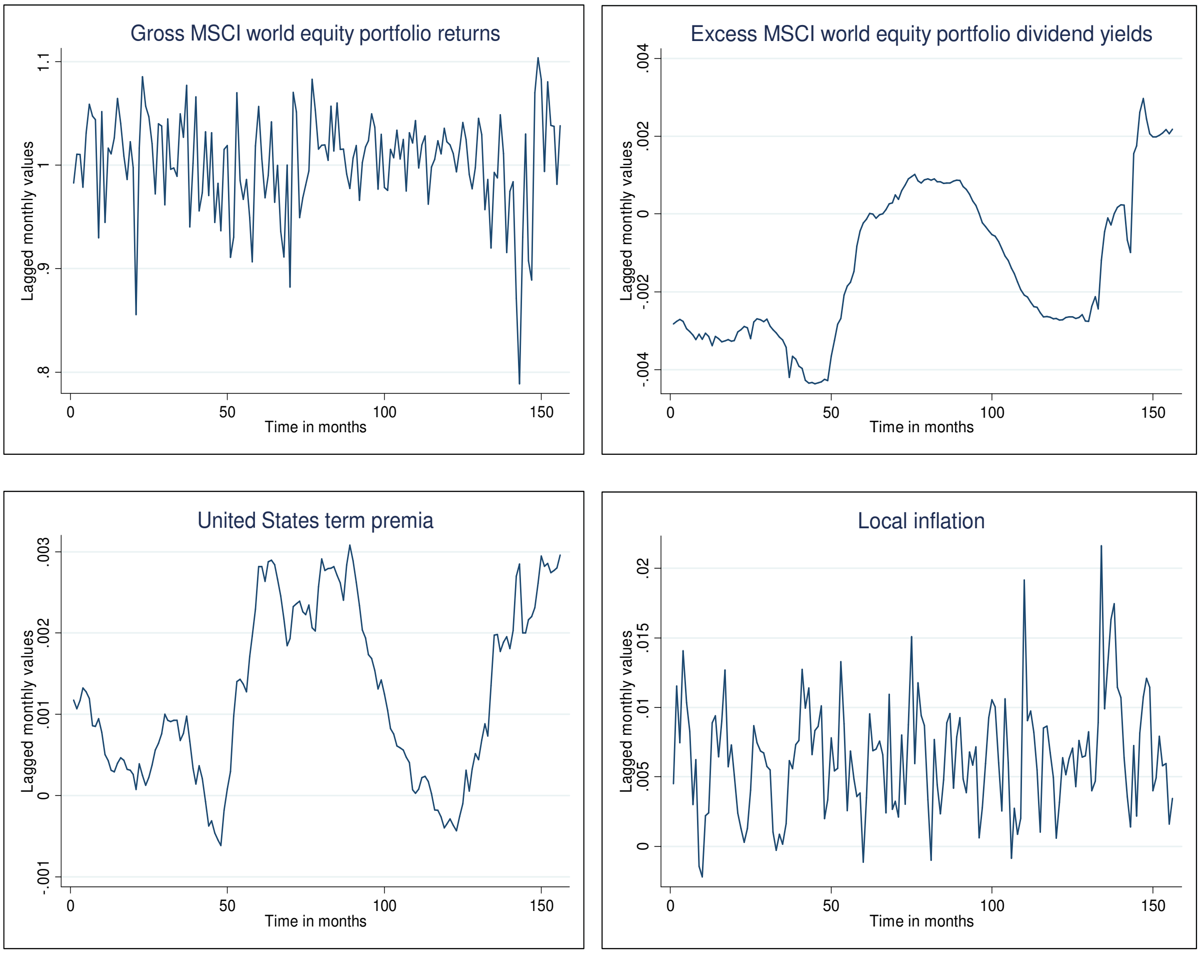


Figure 2 Time series properties of the cyclic components non-stationary instrument variables

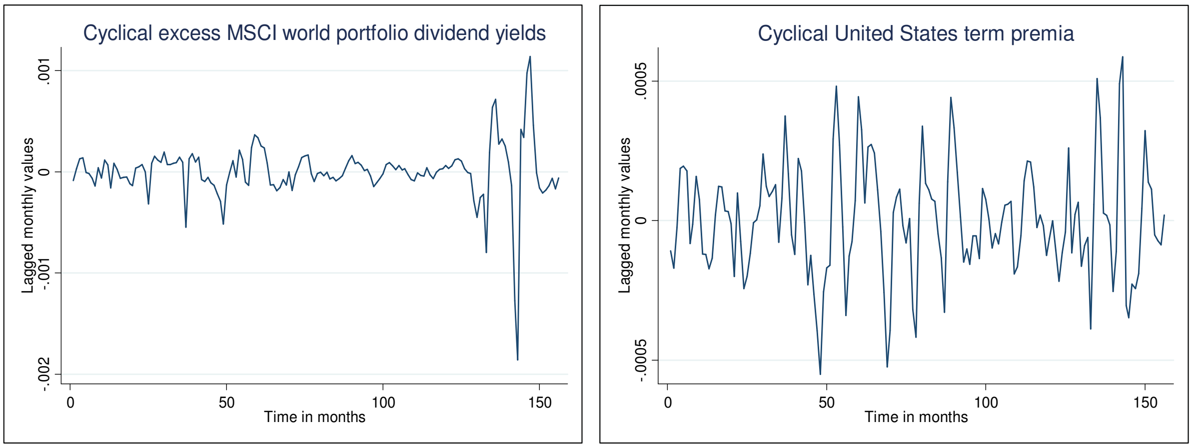
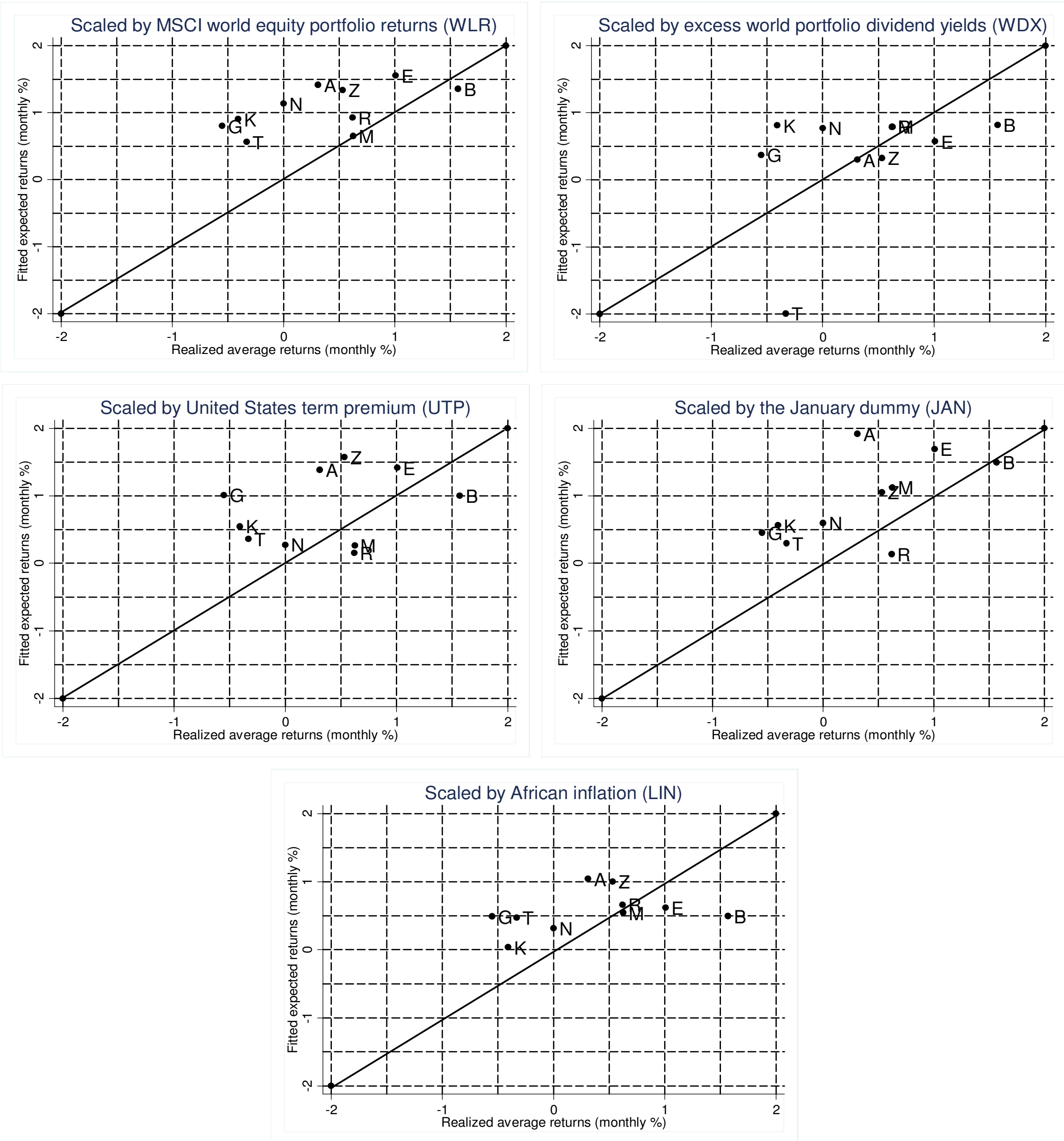
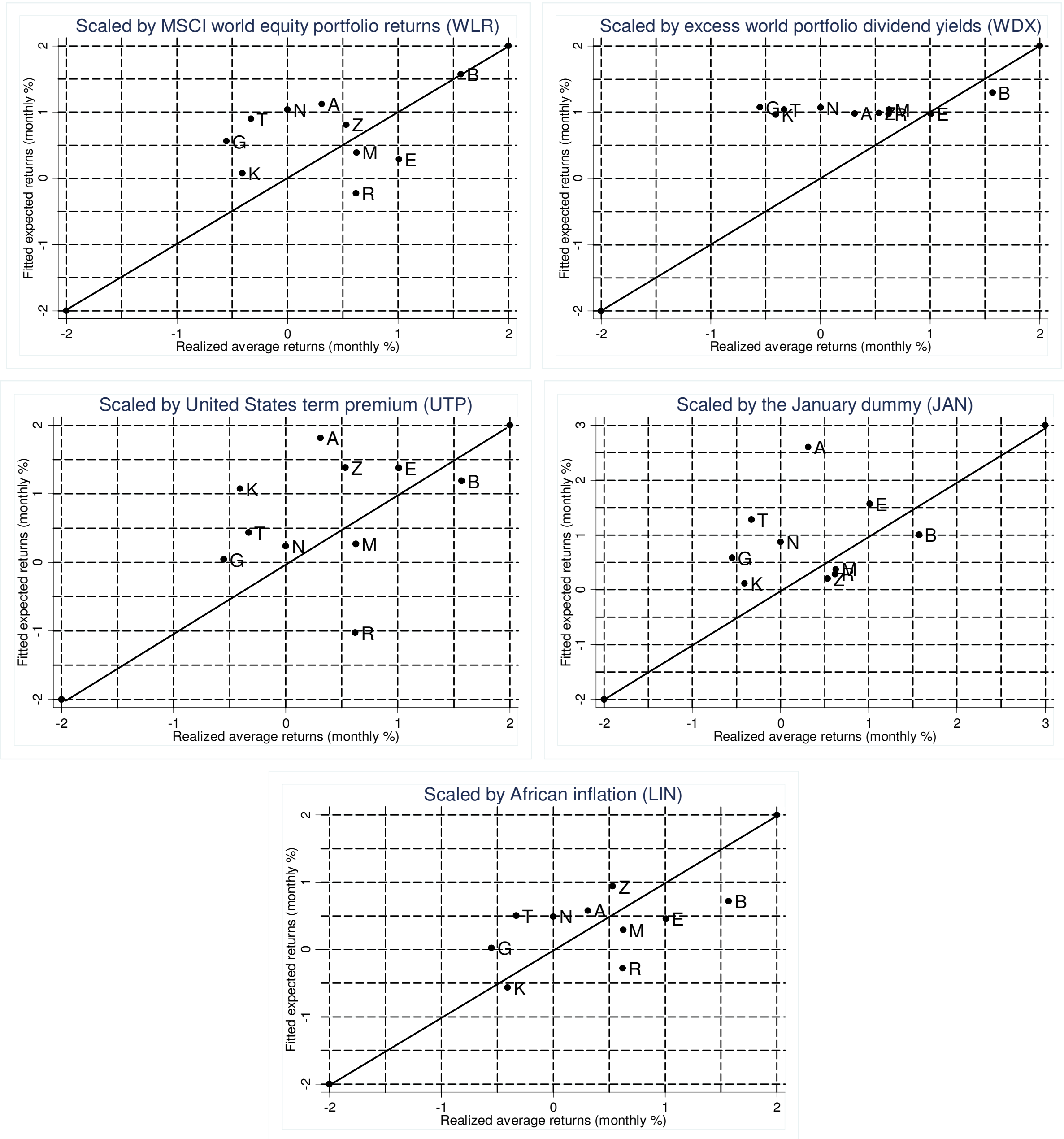


Figure 3 Pricing error plots for the stochastic discount factor model



The figure shows pricing error plots for the stochastic discount factor model of equity returns with MSCI world equity portfolio and real foreign exchange rates as risk factors. Mean realized returns (horizontal axis) are plotted against mean returns implied by the respective model specification (vertical axis). The test assets are equity market aggregate returns for Botswana (B), Egypt (E), Ghana (G), Kenya (K), Mauritius (M) Morocco (R), Namibia (A), Nigeria (N), South Africa (Z) and Tunisia (T). Monthly data cover January 1997 to December 2009. Nominal gross returns denominated in US dollars are used.

Figure 4 Pricing errors plots with nominal exchange rates and inflation as priced factors



The figure shows pricing error plots for the stochastic discount factor model of equity returns with MSCI world equity portfolio nominal foreign exchange rates and an inflation variable as risk factors. Mean realized returns (horizontal axis) are plotted against mean returns implied by the respective model specification (vertical axis). The test assets are equity market aggregate returns for Botswana (B), Egypt (E), Ghana (G), Kenya (K), Mauritius (M) Morocco (R), Namibia (A), Nigeria (N), South Africa (Z) and Tunisia (T). Monthly data cover 1997:01 through 2009:12. Nominal gross returns denominated in US dollars are used.