Financial Globalization & Democracy: Political Uncertainty, Information Acquisition, and Contagion from Domestic to International Investors in Emerging Markets

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Abstract. A central question in international and comparative political economy concerns the constraining effect of global capital markets on emerging-market governments. Global investors vote with their feet by shifting out of a country's assets when uncertainty is high or unfavorable policy changes are anticipated. Despite the increasing globalization of capital, however, I argue that financial market reactions to politics continue to be largely driven by the responses of local investors. Building on models of asymmetric information in capital markets, I posit that better-informed domestic investors form accurate, timely assessments of domestic political risk, while less well-informed international investors observe the signals from domestic investor activity to make their own portfolio decisions. In this model, global market responses to political events in often distant and opaque countries are characterized by contagion dynamics from domestic to foreign investors. I test this contagion model with data on closed-end country funds from 1988-2015, which afford a unique opportunity to measure the differential reactions of international and domestic investors to emerging-market elections. Using a fractional error correction approach, I estimate short- and long-run dependencies between domestic and international investor behavior and find strong evidence of contagion from the former to the latter, confirming that global financial markets transmit political information from locals to foreigners.

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INTRODUCTION

A central question in international and comparative political economy concerns the constraining effect of global capital markets on emerging-market governments. In the past three decades, emerging economies have undergone two consequential, parallel transformations: they have transitioned into democratic regimes and they have become deeply integrated into global capital flows (Santiso, 2013). These transformations have created important political dilemmas for emerging market governments. If, on the one hand, capital-scarce developing nations have greatly benefited from increased access to foreign savings, on the other, financial integration has come at the cost of increased scrutiny over political and policy choices by global investors (Mosley, 2003). Having the ability to vote with their feet by shifting out of a country's assets when unfavorable policy changes are anticipated, global investors have become important political actors in the domestic politics of emerging economies (e.g. Martínez and Santiso, 2003; Campello, 2015).

A crucial step in understanding global market responses to politics in the emerging world involves knowing when and how markets make investment allocation decisions in response to political events. Doing so requires knowledge of how market participants collect and process information about politics, assess country-level risks, and how, in turn, they react to political events in often distant and opaque countries. Indeed, political economists have dedicated considerable effort to inferring investors' preferences from their reactions to changes in country-level political risk. A growing body of cumulative knowledge shows that global capital markets dislike political uncertainty (e.g. Leblang, 2002; Hays, Freeman, and Nesseth, 2003; Frot and Santiso, 2013) and are able to inflict substantial costs on the governments that deviate from investors' preferred policies (Mosley, 2003; Vaaler, Schrage, and Block, 2005; Sattler, 2013).

Despite the increasing globalization of capital, however, I argue that global capital market reactions to politics continue to be largely driven by local dynamics—more specifically, by the responses of local investors. Building on a class of models of asymmetric information in financial markets, I argue that important contagion dynamics arise between domestic and international investors. In this contagion model, better-informed domestic investors are major drivers of global market reactions to changes in country-level political risk. Domestic investors form accurate, timely assessments of domestic political risk and adjust their investment portfolios accordingly. Less well-informed international investors, in contrast, observe the signals from market prices and the trading activity of domestic investors to make their own portfolio allocation decisions. Therefore, the market behavior of domestic investors predicts that of foreign investors, as financial markets transmit information from the former to the latter.

Anecdotal evidence from the financial news pages points to the existence of important differences in the reactions of domestic and foreign investors to potentially destabilizing and uncertainty-inducing political events in emerging countries. In 2002, when polls showed that Lula, the left-wing candidate of the Workers' Party, was a frontrunner in Brazil's presidential elections, pessimism took over capital markets. Fears of a potential debt default, as well as other risks from taxation and regulatory interventions, sent financial markets on a bearish ride. Whereas the pessimistic market reactions to Lula's favoritism in the polls have been widely documented in the literature (Martínez and Santiso, 2003; Jensen and Schmith, 2005; Hardie, 2006), less attention has been paid to the differences in the way domestic and international investors reacted. As investors panicked and capital fled, Lula and his party embarked on an effort to soothe investors' fears and made a series of political gestures intended to regain the market's confidence, including an alliance with a conservative party and the explicit approval of a new IMF loan agreement.[2] As a result, market perceptions of the Lula risk changed gradually. Not all investors, however, updated their beliefs at the same pace. Speaking to a Brazilian newspaper, a local banker noted, "the domestic market has concluded that the [Workers' Party] macro-economic policy is very coherent. Only foreigners had it in their heads that the PT had not changed (...)."[3]

Similarly, in early June 2005, news broke in Brazil about a major corruption scheme involving the purchase of Congressional votes by the governing party. As the political crisis unfolded, Brazilian stock and bond markets plummeted, reflecting doubts about the incumbent government's political fate. However, domestic and foreign investors reacted with different timing. The financial press reported that the first to react were domestic retail investors, known to be especially risk averse when it comes to investing in the stock market.[4] In contrast, foreign investors only seem to have become responsive to the unfolding crisis more than a month after the first news about the scandal came out. According to a major Brazilian newspaper, a Goldman Sachs economist observed that

> up until last week, the extent of foreign investors' worry about the political crisis was very little when compared to the local market. However, foreign investors are now more alert to the crisis, and are trying to understand it better in order to draw scenarios. Up until now, foreigners' bias has been to support the market, while locals have done

[2] "IMF Approves US$30.4 Billion Stand-By Credit for Brazil." *IMF Press Release No. 02/40*. September 6, 2002.

[3] "Lula decries market fears of Brazil debt default." *Financial Times*. October 1, 2002. As cited in Brooks and Mosley (2007).

[4] "Crise espanta pequeno investidor da Bolsa." *Folha de S.Paulo*. August 15, 2005.

the opposite.[5]

Events appear to have followed a similar thread in Mexico's peso crisis of December 1994. Acting on privileged information, local investors led the run on the peso even before the large devaluation was announced. As *The Economist* reported,

> Mr. Zedillo's finance minister pre-announced his devaluation to a coterie of leading Mexican businessmen. (...) The markets concluded that things were out of control. According to bankers and brokers with their ears to the ground, the first to flee were not fickle foreign investors but well-informed Mexicans.[6]

These examples are not isolated. This article offers systematic evidence of the leading role of domestic investors in global market responses to emerging-market politics. I use a fractional error correction approach to examine dynamic interdependencies between domestic and international investors in their reactions to elections in emerging countries. Using daily and weekly data on closed-end country funds that provide a unique opportunity to disentangle the market behavior of local and foreign investors, I present strong evidence of contagion from domestic to international investors. The results show that domestic investors respond immediately to contemporaneous shocks to political risk and transmit these shocks to international investors. Moreover, having higher exposure to domestic risks, local investors tend to respond with a greater rebalancing of their portfolios in the face of variations in political uncertainty. This tendency is reflected in greater abnormal market performance by domestic investors in elections when compared with international investors.

These findings have important implications for our understanding of global capital market responses to politics in emerging markets. They shed light on a particular type of contagion process within international capital markets and highlight how political events in emerging economies shape global capital flows. The conclusions draw attention to the importance of information flows within capital markets in explaining the constraining effects of capital mobility on emerging countries. Furthermore, if foreign and domestic investors operate with different information, use different strategies, and respond in different ways to politics in emerging economies, then it matters which audience governments choose to target when seeking to appease markets. The findings, therefore, reveal a new dimension of the confidence game between markets and governments.

[5] "Analistas já veem 'contaminação' política." *Folha de S.Paulo*. July 26, 2005. See also "Estrangeiros fogem da Bolsa em agosto." *Folha de S.Paulo*. September 5, 2005.

[6] "Another day, another dive." *The Economist*. October 28, 1995.

GLOBAL FINANCIAL MARKETS AND POLITICAL RISK

Financial markets' influence over domestic policy decisions is a longstanding and recurrent question in political economic research. Capital owners have long been thought to hold a privileged political position, since they have at their disposal the threat of exit or defection. By withholding investment decisions, investors can affect the general level of economic activity and the amount of resources available for financing government activity. This ability to punish a country's government for deviations from their preferred policies confers on capital owners an advantage when it comes to defining the contours of government policy (Hirschman, 1978; Lindblom, 1982; Bates and Lien, 1985). Financial globalization has made markets' influence even more salient by increasing capitalists' exit options and by widening the pool of investors that have a stake in emerging markets' domestic policy decisions (Andrews, 1994; Frieden, 1991; Keohane and Milner, 1996; Rodrik, 2000).

Understanding global market responses to domestic politics and policymaking involves deciphering market participants' preferences and uncovering the way they act on those preferences. Political economists have accumulated significant knowledge on what markets value and how they react to domestic political processes. Numerous studies have leveraged variation in political risk during elections to assess investors' sensitivity to expected changes in government policy (Bernhard and Leblang, 2006; Bechtel, 2009; Block and Vaaler, 2004; Vaaler, Schrage, and Block, 2005; Frot and Santiso, 2013; Jensen and Schmith, 2005; Campello, 2015; Leblang, 2002; Hays, Freeman, and Nesseth, 2003; Spanakos and Renno, 2009; Sattler, 2013). In periods of potential political change, investors anticipate future changes in policy and adjust their portfolios accordingly. As they perceive political parties to have distinct agendas and to seek different economic objectives, investor behavior during elections is driven by expectations about how partisan changes in government will affect policy outcomes and, ultimately, investment returns. If market participants expect future changes to adversely affect the value of their assets, they will rebalance their portfolios toward assets that are more insulated from the expected policy change. Indeed, Bernhard and Leblang (2006) show that capital markets systematically react to uncertainty-inducing political processes in a negative way. Frot and Santiso (2013), moreover, note that nearly all major financial crises in emerging economies have coincided with electoral cycles. Similarly, Leblang (2002) finds that elections induce speculative attacks on developing country currencies.

We also know that there is important variation in financial markets' behavior toward national governments. Mosley (2003) shows that global capital markets pressure developed and developing economies to different degrees. Markets exert an intense but narrow influence on advanced economies, while developing

countries seem to suffer broad pressure from sovereign debt holders. Moreover, Campello (2015) and Kaplan (2013) show that global markets' influence on developing countries is greatest during bad economic times—such as periods of low international liquidity—when capital is scarcest and governments' bargaining position is weakest. Considerable variation also exists across different classes of investors. Portfolio and foreign direct investors and—within the class of portfolio investment—equity and sovereign bond investors typically show different degrees of sensitivity to different risk factors (Ahlquist, 2006; Bernhard and Leblang, 2006; Mosley, 2003; Mosley and Singer, 2008; Wellhausen, Forthcoming).

With a few exceptions, political scientists have yet to take seriously the role of information constraints in determining market reactions to politics. How do global investors collect and process costly information about politics in often distant and opaque developing countries? What strategies do they use to deal with these information constraints when assessing country-level risks? In what follows, I posit that domestic and international investors operate in different information environments and that this difference matters for the way they respond to political risk in emerging markets. I argue that differences in information bring about contagion processes from domestic to foreign investors, and that these interdependencies help predict important patterns in the way global financial markets exert pressure on national governments.

DOMESTIC INVESTORS, INTERNATIONAL INVESTORS, AND POLITICAL RISK IN EMERGING ECONOMIES

Domestic and foreign investors differ in two important ways in their responses to domestic political risk in emerging countries. First, building on models of asymmetric information in capital markets, I derive predictions of contagion dynamics from domestic to international investors. Then, I show that because of persistent home bias in global portfolio allocation, we should observe much stronger reactions from domestic than from international investors to shocks in political uncertainty. These two processes combined imply that local investors are major drivers of the reactions of international capital markets to politics.

Information Asymmetry and Financial Contagion from Domestic to International Investors

A long research tradition in finance and macroeconomics has examined the impact of information frictions on price formation in financial markets. This research program shows that the existence of fixed information costs give rise to information asymmetries among market participants. In turn, information asymmetry give

rise to contagion dynamics within financial markets. Grossman and Stiglitz developed the class of seminal information asymmetry models that describe the information transmission role of prices in capital markets (Grossman, 1976; Grossman and Stiglitz, 1976, 1980). Two of the fundamental assumptions of the model are that information is costly and that market participants decide on whether or not to become informed based on the expected utility of acquiring new information. Informed investors, who pay the fixed costs of information, make investment allocation decisions based on fundamentals. In turn, uninformed investors observe the signals from market prices and trading activity to learn from informed investors and choose their own portfolio allocations. Information, therefore, flows from informed to uninformed investors, and the market behavior of informed investors anticipates that of uninformed ones.

When it comes to investment allocation decisions on a global scale, fixed information costs can be pronounced. Investing abroad requires the routine collection and processing of detailed information about the countries involved, such as macroeconomic data and political indicators (Calvo and Mendoza, 2000*b*; Frankel and Schmukler, 1996). Assessing political risk can be particularly challenging, as investors need to take into account political institutions, practices, and the interests of local politicians, businesses, and voters. Furthermore, much of that information cost is fixed—that is, it is independent of country size and the amount of capital involved (Calvo and Mendoza, 2000*c*).

Global investors looking to diversify their portfolios into emerging markets face especially high information costs. Politics is distinctively uncertain in developing countries, since these countries tend to have weakly institutionalized parties and party systems (Mainwaring and Scully, 1995; Dix, 1992; Kuenzi and Lambright, 2001). Party systems in developing countries tend to be more unstable and volatile (Mainwaring and Zoco, 2007; Mozaffar and Scarritt, 2005; Roberts and Wibbels, 1999). Moreover, parties are less cohesive and disciplined (Ames, 1995, 2002), and because party systems tend not to be structured along salient social cleavages (Dix, 1989; Kitschelt et al., 2010), patterns of representation are more personalistic and less programmatic (Keefer, 2007; Kitschelt et al., 2010).

Institutionalization affects uncertainty in different ways: as a source of credible commitments and as a source of information about politicians' ideological and programmatic preferences. As a commitment device, party organization enforces coordination among politicians and creates credible commitments of parties to policy positions (Carey, 2007; Kitschelt et al., 2010; Lupu and Riedl, 2013). As an informational device, institutionalized party systems reduce information costs by making party affiliation informative of politicians' ideologies and policy preferences (Kitschelt et al., 2010; Mainwaring and Torcal, 2006; Mainwaring and Zoco, 2007). Weak institutionalization, therefore, increases political risk and, more im-

portantly, makes it costlier for investors to form accurate assessments of that risk.

Moreover, information costs should be especially salient for foreign investors when compared to domestic investors. Proximity and familiarity with the local political environment give local investors an advantage in terms of information costs. As Obstfeld (1998) points out, "unfamiliarity with foreign products, firms, business practices, accounting standards, political trends, and regulatory environments" plays a role in investors' decisions to diversify towards foreign markets. Domestic investors often have more favorable access to locally available information, and foreign investors may need to incur extra costs to obtain the same information (Frankel and Schmukler, 1996, 2000).

If domestic investors have a cost advantage when it comes to information, they should be able to take positions in the market that are better than those of international investors. Indeed, research in finance shows that location matters for investment performance. Choe, Kho, and Stulz (2005) show that foreign investors obtain lower returns in the Korea stock exchange. Chang (2010) uses an interesting research design to measure the effects of location on investment performance. He compares expatriate firms established in emerging markets to similar foreign investment firms operating in the same market from abroad and finds that physical presence in the economy matters. Similarly, Coval and Moskowitz (2001) report that fund managers earn substantial abnormal returns in geographically proximate investments, and Teo (2009) shows that hedge funds that have a physical presence in their main investment region outperform hedge funds with no local presence. Moreover, Hau (2001) highlights information costs from language barriers. He finds that foreign traders in non-German-speaking financial centers earn inferior profits in their trading of German stocks than foreign traders in German-speaking centers. Finally, using a cross-country design and a sample of thirty-two countries, Bae, Stulz, and Tan (2008) show that local investment analysts make better earnings forecasts for home stocks than foreign analysts do. The local advantage is greater in countries where firms are less transparent, that is, where information costs are higher. This is consistent with the expectation that domestic investors should have a higher advantage in weakly institutionalized polities, where political information is costlier.

While differences in information costs can bring about information asymmetries between foreign and domestic investors, another set of incentives should reinforce these asymmetries. As Calvo and Mendoza (2000*a*) show, financial globalization reduces incentives for information gathering, since the expected gains accrued by paying fixed information costs tend to fall as the number of countries where the wealth can be invested grows. This occurs because, as the number of countries in a global portfolio increases, the contribution of each country to the portfolio risk decreases, thereby reducing the marginal return of information on

any given country. However, home bias in portfolio allocation implies that this decrease in the incentives to collect information will affect domestic and foreign investors differently. Home bias means that investors tend to hold a larger fraction of their investment portfolios in the form of domestic securities than would be optimal under standard portfolio theory (French and Poterba, 1991; Tesar and Werner, 1995; Lewis, 1999). It also implies that, for any given country, local investors will tend to hold a greater share of their portfolios in that country's assets than foreign investors will. Under home bias, local investors' interests should be more closely tied to the performance of the local market. As a result, domestic investors should derive higher gains from collecting and processing local information. In the presence of home bias, financial globalization reduces overall incentives for information gathering, but increases the incentive gap between domestic and foreign investors, thus reinforcing information asymmetries between the two groups.

Domestic investors, therefore, will tend to be on the informed side of the market more often than international investors when it comes to investing in emerging markets. But how do international investors cope with information asymmetry? Existing research suggests that investors often devise cost-efficient strategies to deal with information limitations, and that these strategies have implications for how markets respond to policy and political change (Brooks, Cunha, and Mosley, 2015; Mosley, 2003). When it comes to foreign-domestic asymmetries, an efficient strategy is to extract signals from the market itself (Grossman, 1976; Calvo and Mendoza, 2000*c*; Frankel and Schmukler, 2000). Under conditions of costly information, informed domestic investors will bid asset prices up or down based on acquired information on fundamentals; less well-informed international investors will observe signals from market prices and trading activity and learn about fundamentals from informed investors. Therefore, under information asymmetry, information should flow from informed to uninformed investors, and the market behavior of informed investors should predict that of uninformed ones. This should create dynamics of contagion from domestic to international investors, whereby better informed domestic investors react immediately to changes in country-level risks, and international investors react with a delay after observing the market behavior of their domestic counterparts.

Differential Sensitivity to Country-Level Risks

The second process that brings about differences in the way domestic and international investors respond to shocks to country-level risks is differential risk exposure. Here, too, home bias in global portfolio allocation plays an important role. As already discussed, home bias refers to an observed pattern in international

finance according to which the proportion of foreign assets held by domestic investors is too small relative to the optimal predictions of standard portfolio theory. In other words, home bias implies that even with the expansion of diversification opportunities brought about by the advance of financial globalization, investors remain disproportionately attached to their local markets.

Home bias is a stylized fact of international finance and an open empirical question. French and Poterba (1991) report that in 1989 U.S. and U.K. investors held around 8% of their portfolio in foreign shares, with investors in other advanced economies holding a similar proportion of foreign equities. In 2000, ownership of foreign equity had risen to only 12% of the U.S. portfolio (Warnock, 2002). Moreover, Bekaert and Wang (2009) show that home bias is particularly pronounced among investors from emerging markets. Therefore, large gains from international diversification remain unrealized and investors remain disproportionately exposed to domestic risks.

I take home bias as a given and treat it as an exogenous explanatory factor. The consequences of home bias for global market reactions to elections and other country-level variations in political risk are straightforward. Being more exposed to domestic risks, local investors should rebalance their portfolios to a greater extent than international investors in the face of increases or decreases in country risk. When it comes to election-induced portfolio adjustments, we should expect stronger responses from domestic investors. This should be the case both under predictable and unpredictable elections. When market actors can easily forecast election outcomes, we should expect them to anticipate any future changes in the orientation of economic policy and adjust their portfolios before those changes take place. If investors anticipate that a new government will adopt policies that will decrease the return to their assets, they will shift their portfolios out of vulnerable assets towards safer ones. As result, the price of those assets will fall. In contrast, an expected change in policy with favorable impact on returns should have the opposite effect. Because an unfavorable turn in economic policy affects domestic investors' portfolios disproportionately, they should shift a larger share of their portfolio towards safer assets when compared to international investors.

The same is true for unpredictable elections. Unpredictable elections cause uncertainty over the future course of economic policy and thus increase the risk of assets that are vulnerable to potential policy changes. When the resolution of the uncertainty surrounding and event takes time to resolve, investors demand higher returns for holding the risky asset and thus drive prices down before that uncertainty is resolved (Brown, Harlow, and Tinic, 1988; Block and Vaaler, 2004; Bernhard and Leblang, 2006). In other words, when election outcomes remain difficult to predict until the very end of the process, investors with a stake in the election results will drive down the price of assets that are vulnerable to the election

result. Therefore, asset returns should be negatively affected by electoral uncertainty. Moreover, election-induced increases in portfolio risk should be proportional to the participation of vulnerable assets in the portfolio. Domestic investors, who hold a larger share of their portfolio in local assets than foreign investors do, should see a sharper increase in their overall portfolio risk, and thus should more aggressively hedge against those risks. In sum, we should observe larger responses to political risk on the part of domestic investors than on the part of foreign investors.

EMPIRICAL STRATEGY

Measuring Foreign and Domestic Investor Behavior: Closed-End Country Funds

In this section, I present an empirical strategy to measure and model contagion dynamics between foreign and local investors, as well as test differences in the way these two classes of investors respond to electoral uncertainty in emerging economies. To investigate differences in foreign and domestic investor responses to politics, it is necessary to disaggregate market transactions by country of domicile. However, typical financial data—equity and bond prices, interest rate spreads on sovereign debt—are aggregate market data, and thus do not allow for the distinction between foreigners and nationals.

Telling apart the reactions of foreign and domestic investors requires a measure that can capture the differential market sentiments of each group. Data from closed-end country funds provide such a measure. Closed-end country funds are publicly traded investment companies that trade on the open market and hold and manage portfolios concentrating in equity markets of particular countries. A closed-end country fund—henceforth country fund—consists of a fixed number of shares that are invested in a set of stocks from a particular country. Unlike open-end funds—such as conventional mutual funds—once a closed-end fund is established new shares cannot be issued and existing shares cannot be redeemed. Investors desiring to buy or sell country-fund shares need to trade them on secondary markets as regular stocks.

Country funds are traded in several developed-country security markets and allow investors to diversify their portfolios internationally without needing specific knowledge of particular industries or firms (Frankel, 1995; Hardouvelis, La Porta, and Wizman, 1995; Frankel and Schmukler, 1996, 2000). In the New York Stock Exchange, for example, they trade at their U.S. dollar price. A fund's net asset value—NAV—is defined as the per-share dollar value of the fund's underlying assets; that is, the NAV is the aggregate value of the fund's constituent equities. The country fund premium is the difference between the fund price and its NAV. In a perfectly efficient and frictionless market, country fund premia would be zero,

since the fund price and NAV are essentially two market values of the same asset (Stulz, 1981). In practice, fund premia and discounts (negative premia) can be large and vary widely over time and across countries.

For the purposes of this study, country fund prices and NAVs provide a useful measure of the dependent variable—the market behavior of international and domestic investors, respectively—as they reflect how country fund holders value their assets relative to holders of the individual shares and, therefore, capture differences in the sentiment of foreign and domestic investors towards a particular country (Frankel and Schmukler, 1996, 2000). Frankel and Schmukler (1996, 512), for example, argue that the price of the country fund, which is traded in international markets, reflects relatively better the information and expectations held by international investors, while the NAV, which is determined in local stock markets, reflects relatively better the information and expectations held by local investors. In their words, "the discount or premium on a country fund becomes a measure of the spontaneous pessimism or optimism with which U.S. investors view the country in question, relative to the investors within that same country" (Frankel, 1995, 17). Therefore, a premium indicates that foreign investors have more favorable expectations (relative to domestic investors) regarding a country's economic prospects; a discount indicates that domestic investors have relatively more favorable expectations. In sum, the existence of country fund premia and discounts provide a unique opportunity to investigate and contrast the behavior of foreign and domestic investors towards political events in emerging markets.

Table 1 shows the sample of twenty-one closed-end funds focused on individual emerging economies that are used in the analysis. The sample has wide geographic coverage, including funds with a focus on countries in Africa, Asia, and Latin America. The temporal coverage is also extensive and varies by fund: weekly data start as early as 1988 and go up to 2015, while data with daily frequency are available for fewer funds and range from 1992 to 2015. Country fund premia show considerable variation over time and across countries. Figure 1 shows the time series of weekly country fund premia/discounts for the included funds. As shown in Figure A1 in the Data Appendix, emerging-market country funds typically trade at a discount. However, there is large variation, and funds also trade at a premium over several periods covered by the data.

Modeling Contagion from Domestic to International Investors: A Fractional Error Correction Approach

The information asymmetry model predicts a process of contagion from domestic to international investors. Because local investors face lower costs to acquiring domestic political information, they are expected to track domestic develop-

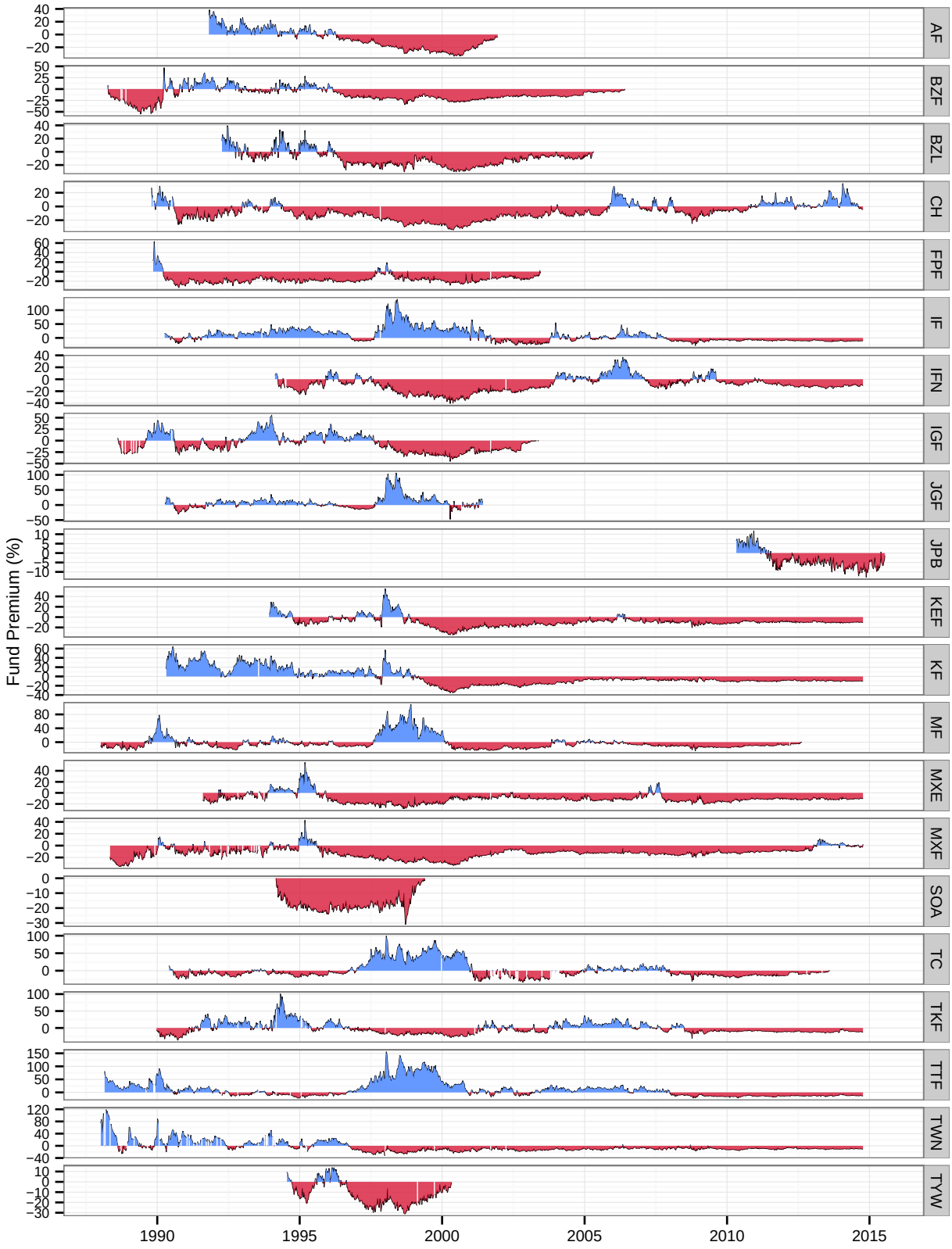

Figure 1. Closed-end emerging-market country fund premia.
Note: Blue-shaded areas represent a positive premium and red-shaded areas represent a negative premium (discount). For an explanation of country fund ticker symbols, see Table 1.

Table 1. Sample of emerging-market closed-end country funds (weekly and daily data).

		Weekly frequency		*Daily frequency*	
Fund	*Symbol*	*N*	*Period*	*N*	*Period*
Argentina Fund	AF	530	10/25/1991 - 12/14/2001	1042	12/18/1997 - 12/14/2001
Brazilian Equity Fund	BZL	682	04/10/1992 - 04/29/2005	1861	03/13/1998 - 04/29/2005
Brazil Fund	BZF	949	04/08/1988 - 06/09/2006	2207	12/18/1997 - 06/02/2006
JPMorgan Brazil Investment Trust	JPB	273	04/30/2010 - 07/17/2015	1367	04/26/2010 - 07/21/2015
Chile Fund	CH	1304	10/20/1989 - 10/10/2014	4094	02/03/1999 - 10/13/2014
First Philippine Fund	FPF	711	11/10/1989 - 06/20/2003		
India Fund	IFN	1078	02/18/1994 - 10/10/2014	1292	10/30/2009 - 10/13/2014
India Growth Fund	IGF	772	08/12/1988 - 05/23/2003		
Indonesia Fund	IF	1280	04/06/1990 - 10/10/2014	4093	02/04/1999 - 10/13/2014
Jakarta Growth Fund	JGF	583	04/12/1990 - 06/08/2001	861	02/20/1998 - 06/08/2001
Korea Equity Fund	KEF	1089	12/03/1993 - 10/10/2014	3552	03/02/2001 - 10/13/2014
Korea Fund	KF	1277	04/27/1990 - 10/10/2014	4638	01/02/1997 - 10/13/2014
Malaysia Fund	MF	1285	01/08/1988 - 08/17/2012	5367	01/28/1992 - 08/22/2012
Mexico Equity & Income Fund	MXE	1210	08/09/1991 - 10/10/2014	3410	09/17/2001 - 10/10/2014
Mexico Fund	MXF	1380	05/06/1988 - 10/10/2014	4944	11/01/1995 - 10/13/2014
New South Africa Fund	SOA	275	03/04/1994 - 06/02/1999		
Taiwan Equity Fund	TYW	303	07/22/1994 - 05/05/2000		
Taiwan Fund	TWN	1397	01/08/1988 - 10/10/2014	2687	06/25/2004 - 10/13/2014
Thai Capital Fund	TC	1212	06/01/1990 - 08/16/2013		
Thai Fund	TTF	1390	02/26/1988 - 10/10/2014		
Turkish Investment Fund	TKF	1295	12/22/1989 - 10/10/2014	3552	03/02/2001 - 10/13/2014

Sources: Bloomberg and Lipper/Thomson Reuters.

ments more closely and to be the first to adjust their portfolios to new information. According to this contagion model, short-run changes in NAVs, which contain more information about local conditions, should help predict short-run changes in prices, while changes in prices should not be as good a predictor of changes in NAVs. More importantly, we should expect domestic and foreign investors to be in a long-run equilibrium relationship. Fund prices and NAVs are two prices of the same underlying basket of assets, so in the long-run they should reflect fundamentals and not veer away from each other for too long. As new information comes along, domestic investors immediately adjust their portfolios, thus moving the NAV away from equilibrium. As international investors observe signals from local investor activity, they adjust their portfolios accordingly and bring the fund price closer to the NAV, reestablishing the equilibrium. Country fund prices, therefore, should be the variable to adjust to deviations from the long-run equilibrium.

An error correction approach is a natural candidate for a modeling strategy.

Vector error correction models allow us to model the short- and long-run relationship between two or more series. Short-run changes in each series are modeled as a function of past changes in the series itself as well as past changes in every other series in the system. Moreover, each series is also modeled as a function of past deviations from the long-run equilibrium.

The modeling of responses to deviations from a long-run equilibrium rests on the notion of cointegration. Two non-stationary, or integrated, series are said to be cointegrated if a linear combination of them is stationary, or non-integrated. Stationarity implies that a series has a constant mean and finite variance. Stationary, or $I(0)$, series have short memory and are mean reverting. In contrast, non-stationary, or $I(1)$, series have infinite memory; an exogenous shock to the series permanently sends it away from the previous mean (Box-Steffensmeier et al., 2014). Financial asset prices, such as country fund prices, typically follow a random walk process and are thus non-stationary (for a review of the extensive literature on the predictability of asset prices, see Cochrane, 2005, 389-454). Frankel and Schmukler (2000), for example, find evidence that country fund NAVs and prices are $I(1)$.

The determination of the order of integration is important because statistical regressions of integrated series in level form run a high risk of spuriousness (Hamilton, 1994). A common solution is to estimate a time series model in differences. Indeed, financial asset prices are typically differenced and analyzed in the form of asset returns (Campbell, Lo, and MacKinlay, 1997). However, if the variables in question share a long-run relation, an analysis in differences will be misspecified, since the level of the series will contain information that is not present in the short-run changes of the series alone (Engle and Granger, 1987; Hamilton, 1994). The notion of cointegration captures these long-run dynamics by positing that a set of non-stationary series can be tied together by a long-run equilibrium relation. Cointegration has been documented for various financial series, such as exchange rates (Baillie and Bollerslev, 1989), bond spreads (Campbell and Shiller, 1987), and stock prices and dividends (Diba and Grossman, 1988). When it comes to closed-end country funds, however, the evidence for cointegration is mixed. Frankel and Schmukler (2000) find that NAVs and fund prices in Europe, Latin America, and Asia are in a long-run equilibrium relation with a slow rate of adjustment. Levy-Yeyati and Ubide (2000), on the other hand, find that emerging market country fund discounts are nonstationary, suggesting that prices and NAVs do not display equilibrating behavior. And Chang, Eun, and Kolodny (1995) find prices and NAVs to be cointegrated for the majority of country funds from North America and Europe, but not for those investing in Asia.

I adopt a more general approach to cointegration and allow for fractional cointegration in country fund prices and NAVs. Fractional cointegration relaxes the

assumption that the long-run equilibrium between two series is stationary and allows its order of integration to be $I(d)$, with $0 < d < 1$ (Baillie and Bollerslev, 1994; Box-Steffensmeier et al., 2014; Cheung and Lai, 1993; Dueker and Startz, 1998). As such, fractional cointegration provides a broader characterization of the long-run equilibrium relation between country fund prices and NAVs. Whereas in traditional cointegration deviations from equilibrium must be quickly corrected, fractional cointegration is both long memoried and mean reverting. Shocks to fractionally cointegrated processes are persistent in the short-run but dissipate over a longer horizon. Therefore, fractional cointegration allows for the possibility that deviations from equilibrium are highly persistent. A number of financial and economic processes have been found to be fractionally cointegrated, including exchange rates (Baillie and Bollerslev, 1994), stock prices and dividends (Caporale and Gil-Alana, 2014), and spot and future stock indices (Lien and Tse, 1999). Importantly, the mixed evidence of cointegration in country fund prices and NAVs suggests the possibility of fractional cointegration, since traditional unit root tests have low power against fractional alternatives (Diebold and Rudebusch, 1991; Hassler and Wolters, 1994).

Therefore, I model the short- and long-run dependencies between country fund prices and NAVs using a fractional error correction model (FECM) that allows for a long-memoried equilibrium relation between the two series. We can represent the bivariate FECM as follows (Granger, 1986; Dittmann, 2004):

$$\Delta P_t = \phi_1 + \alpha_1[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L} \gamma_{1i}\Delta N_{t-i} + \sum_{i=1}^{L} \omega_{1i}\Delta P_{t-i} + \upsilon_{1t},$$

$$\Delta N_t = \phi_2 + \alpha_2[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L} \gamma_{2i}\Delta P_{t-i} + \sum_{i=1}^{L} \omega_{2i}\Delta N_{t-i} + \upsilon_{2t},$$

where P_t is the log of the country fund price, N_t is the log of the NAV, ϕ_1 and ϕ_2 are intercept terms, ω_i are the coefficients on lagged changes in the series themselves, γ_i are the coefficients on the lagged changes in the other series, and υ_{1t} and υ_{2t} are white-noise error terms. $\alpha[(1-B)^d - (1-B)]z_t$ is the fractionally integrated error correction term, in which B is the backshift (or lag) operator, d is the fractional differencing parameter, and z_t are the equilibrium errors from the cointegration equation, as defined below. Moreover, α_1 and α_2 are the speed-of-adjustment parameters; they measure how fast the series return to equilibrium after a shock that sends them apart. When $d = 0$, the model reduces to a conventional error correction model, in which only the lagged equilibrium error enters the equation.

I add exogenous market risk factors known to influence returns on fund prices and NAVs: the return on a world stock market index (MSCI World), the return on an emerging market stock index (MSCI Emerging Markets), the return on an index

of large U.S. stocks (S&P 500), and the return on an index of small-capitalization U.S. stocks (Russell 2000) (Hardouvelis, La Porta, and Wizman, 1995; Levy-Yeyati and Ubide, 2000). Controlling for these market influences ensures that model results reflect variation in country-specific risks net of global market movements.

For estimation of the FECM, I adapt the Engle and Granger (1987) two-step procedure to the fractional cointegration context. I begin by estimating the fund price and NAV long-run, cointegrating relationship, $P_t = \lambda_0 + \lambda_1 N_t + z_t$, using ordinary least squares and then use the cointegrating residuals, $\hat{z}_t$, as an estimate of the equilibrium errors. Next, I determine the order of integration of fund prices, NAVs, and the cointegrating residuals. I start with unit root tests—the Augmented Dickey-Fuller, KPSS, and Lo-MacKinlay variance-ratio tests—to ascertain whether any of the series are $I(1)$. I proceed to obtain numerical estimates of the fractional integration parameter, d, for each of the series. I use Sowell (1992)'s maximum likelihood estimator as the primary estimator for d, but also obtain estimates from the Geweke and Porter-Hudak (1983), Reisen (1994), and Robinson (1995) estimators to ensure robustness of the results. Additionally, I compute confidence intervals to test in each case whether $d = 0$ and $d = 1$. For each country fund, if price and NAV have the same order of integration, and if their order of integration is greater than that of the cointegrating residuals, then we can conclude that price and NAV are fractionally cointegrated.

In those cases in which the cointegrating residuals are fractionally integrated, I use the maximum likelihood estimate of d to fractionally difference the series. The resulting series can then be used as an estimate of $[(1-B)^d - (1-B)]z_t$ for obtaining estimates of the FECM parameters. I estimate the bivariate system using a vector autoregression of prices and NAVs in differences for each of the country funds in the sample. The lag order for the short-run component in each estimation is determined using Schwarz's Bayesian information criterion (BIC) (Lütkepohl, 1985). The same estimation procedure is followed for both daily and weekly data.

With estimates from the fractional error correction model, we can test substantive hypotheses about contagion from domestic to international investors. The information asymmetry model posits that short-run changes in NAVs, which contain more information about domestic events, should help predict short-run changes in prices, while changes in prices should not be as good a predictor of changes in NAVs. We can test the hypothesis of short-run adjustment with an F-test for the joint hypothesis that all γ_i are zero. From the contagion model, we expect to reject the null hypothesis for γ_{1i} and to fail to reject the null for γ_{2i}. More importantly, we can test whether and to what extent fund prices and NAVs respond to deviations from equilibrium by assessing the statistical significance and relative magnitude of the estimated parameters on the error correction term, $\hat{\alpha}_1$ and $\hat{\alpha}_2$. I expect $\hat{\alpha}_1$, the rate of adjustment of prices, to be statistically significant and large, since inter-

national investors are expected to observe signals from the NAV and adjust their portfolios accordingly. In the same vein, since domestic investors are expected to lead the response to domestic events, there is no reason to expect them to display any equilibrating behavior. Therefore, I expect $\hat{\alpha}_2$ to be statistically insignificant.

Assessing Investors' Reactions to Elections

Not only should information flow from domestic to foreign investors, but, given differences in exposure to country risks, domestic and international investors should also show differences in their sensitivity to electoral uncertainty. Because local investors have greater exposure to local political risk, they should promote larger adjustments to their portfolios than international investors in the face of high political uncertainty. In other words, domestic investors should be more responsive to electoral uncertainty in emerging markets.

I use different strategies to test this claim. I start with a baseline test of investors' reactions to elections by adding an exogenous dummy variable, E_t, to the weekly FECMs coded one for the weeks preceding an election and zero otherwise. I use data from Bormann and Golder (2013) on election dates and code pre-election windows of one, two, and three months, based on the typical length of electoral campaign periods identified in the literature (Block and Vaaler, 2004; Vaaler, Schrage, and Block, 2005; Bernhard and Leblang, 2006). This measure captures the average abnormal return to fund prices and NAVs during electoral periods relative to the normal return in non-electoral periods. If, on average, domestic investors are indeed more responsive to electoral uncertainty and anticipated changes in policy, then we should be able to detect changes in the average return during electoral periods more often for country fund prices than for NAVs. It also follows that the magnitude of changes in returns should be greater in the case of NAVs.

This design, however, has low power to detect market reactions to elections since it does not explicitly account for heterogeneity in the degree of uncertainty across elections and in the direction of expected changes in policy or government partisanship. On the one hand, close elections should induce higher risk-hedging activity by financial investors; for that reason, there should be stronger market reactions in such cases (Campello, 2007). On the other hand, differences in the expected direction of changes in government partisanship should be reflected in differences in the direction of portfolio adjustments. Previous work has shown that markets tend to react negatively to changes towards the left, but to react positively or not at all when a change towards the right or no change is expected (Spanakos and Renno, 2009; Campello, 2015). Data constraints, however, hinder the direct incorporation of changes in partisanship into the analysis. The number of elections

in the sample that have seen partisanship changes from the right or center towards the left is too small. As such, there is not enough variation in the measure to allow for estimation of the FECM.

To address this issue, I leverage the availability of country fund data with daily frequency and combine the fractional error correction approach with an event study design. Event studies are frequently used by financial economists to measure the effect of an event on financial asset prices (Campbell, Lo, and MacKinlay, 1997). An event study of the effect of elections on country fund prices and NAVs involves measuring the abnormal price and NAV return during the pre-election window and comparing it to the normal return in the non-electoral periods. The abnormal return is the actual return of the fund over the election window net of the normal return of the fund over that same window. The normal return is defined counterfactually as the return that would be expected if the election did not take place. In the context of a multivariate error correction model, we can measure abnormal returns during elections by adding a dummy variable for each day in the election window (Binder, 1985; Campbell, Lo, and MacKinlay, 1997). Therefore, I estimate daily FECMs with the added exogenous term,

$$\sum_{\tau=T_1}^{T_2} \beta_\tau D_{\tau,t},$$

where $D_{\tau,t}$ are dummy variables coded one for pre-election day $t = \tau$ and zero otherwise, for $\tau = T_1, T_1 + 1, \ldots, T_2$. The election window is thus defined as $T_1 \leq \tau \leq T_2$, where T_2 is the election day and T_1 is set at either 30, 60, or 90 days prior to T_2. Estimates of β_τ provide us with daily abnormal returns in the pre-election period. We can then aggregate them into cumulative abnormal returns (CAR) by summing $\hat{\beta}_\tau$ over the entire election window.

I estimate a FECM with an event study specification for every general election occurring in the period covered by the daily country fund data (see Table 1 for the temporal coverage). Cumulative abnormal return estimates provide a measure of the cumulative effect of electoral uncertainty on country fund prices and NAVs in each election. If, as hypothesized, domestic investors are more sensitive to domestic electoral uncertainty than foreign investors, then we should observe larger $\widehat{CAR}$s for NAVs than for fund prices. We can verify whether or not this is the case by testing the hypothesis that $\sum \beta_\tau = 0$ with a χ^2 test, as well as by comparing the magnitude of $\widehat{CAR}$ for prices and NAVs.

This design has the advantage of avoiding the difficulties caused by unobserved heterogeneity across elections discussed above. While it does not explicitly incorporate election characteristics—such as the degree of predictability and the direction of the expected change in government partisanship—it does allow us to focus

on the differential reactions of domestic and foreign investors by estimating separate FECMs by election and thus holding unobserved factors constant in each case.

Finally, I also use a strategy that aims to capture the dynamic acquisition of political information by investors during elections. I leverage the availability of high frequency country fund data to assess whether and how foreign and domestic investors update their expectations of potential political change and act upon these expectations. Using data on presidential polls for a number of elections, I test whether investors react to the release of new information about the predictability of election outcomes during election years. I use Freeman, Hays, and Stix (2000)'s measure of entropy, which measures political uncertainty according to the formula: $Entropy_t = 1 - 4[(p_t - 0.5)^2]$, where p_t is the frontrunner's share of the two-party vote. The measure has a maximum value of 1 when $p = .5$ and a minimum value of 0 when either $p = 0$ or $p = 1$. Entropy measures uncertainty as being highest when the two leading candidates have equal probability of winning; in contrast, entropy is lowest when election outcomes are most predictable. I compute entropy measures using polling data from Jennings and Wlezien (Forthcoming) for those cases in which both polling and daily country fund data are available. In addition, I use polling data for the 2002 Brazilian election from Jensen and Schmith (2005), and collect additional data for the 1998 and 2014 Brazilian elections. I add entropy as an exogenous term to the FECM and estimate separate models for each election. Each estimated model covers a period of up to one year preceding the election, with the actual coverage in each case depending on the length of the available polling series.

RESULTS

The contagion model based on asymmetric information between domestic and international investors predicts specific patterns of dynamic interdependence between the market behavior of those two classes of financial investors. An examination of the dynamic properties of closed-end country fund prices and net asset values (NAVs) shows that many of the predictions are borne out in the data. I begin by examining estimates for the order of integration and cointegration of fund prices and NAVs. Results of the unit root tests are reported in Tables A1-A6 of the Data Appendix. Using both the weekly and the daily series, the Augmented Dickey-Fuller test cannot reject the null of a unit root for either fund prices or NAVs for every fund in the sample, while the KPSS test rejects the null of stationarity in all cases. The Lo-MacKinlay variance-ratio test, however, rejects the null of a random walk with drift for both prices and NAVs for all funds. The battery of unit root tests, therefore, points towards an $I(1)$ diagnostic for prices and NAVs,

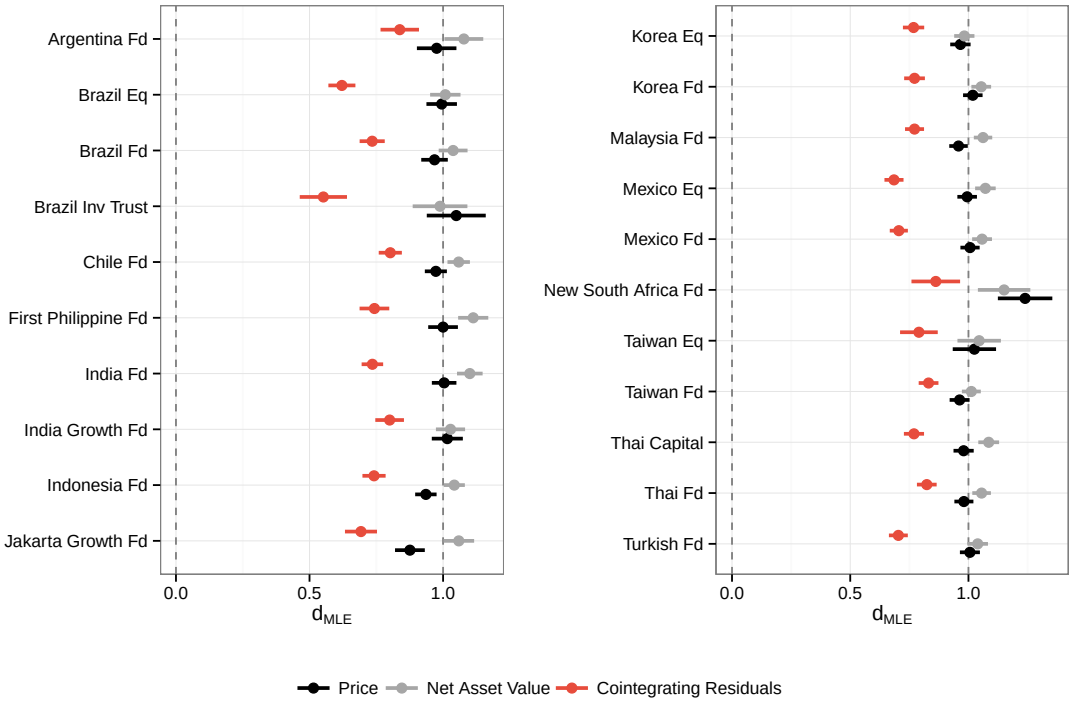

Figure 2. Fractional cointegration of country fund prices and net asset values. Weekly country fund data. Maximum likelihood estimates of the fractional differencing parameter, d. Horizontal bars are 95% asymptotic confidence intervals.

but is inconclusive.

I also run the same unit root tests on the residuals of the price and NAV cointegration equations. For almost every fund, the Augmented DF test and the variance-ratio test reject the null of a random walk, while the KPSS test rejects the null of stationarity. We cannot, therefore, conclude that prices and NAVs are cointegrated in the conventional way, since deviations from equilibrium are not stationary. However, these results strongly suggest the possibility of fractional cointegration.

Indeed, estimates of the order of fractional integration, $I(d)$, for country fund prices, NAVs, and their cointegrating residuals confirm that prices and NAVs are fractionally cointegrated, that is, they are tied together in a long-run equilibrium relation characterized by long memory. Figures 2 and 3 show maximum likelihood estimates of the d parameter for the weekly and daily country fund series. First, for almost all funds the estimated d of prices and NAVs cannot be statistically distinguished from one, suggesting that these series are in fact integrated. In

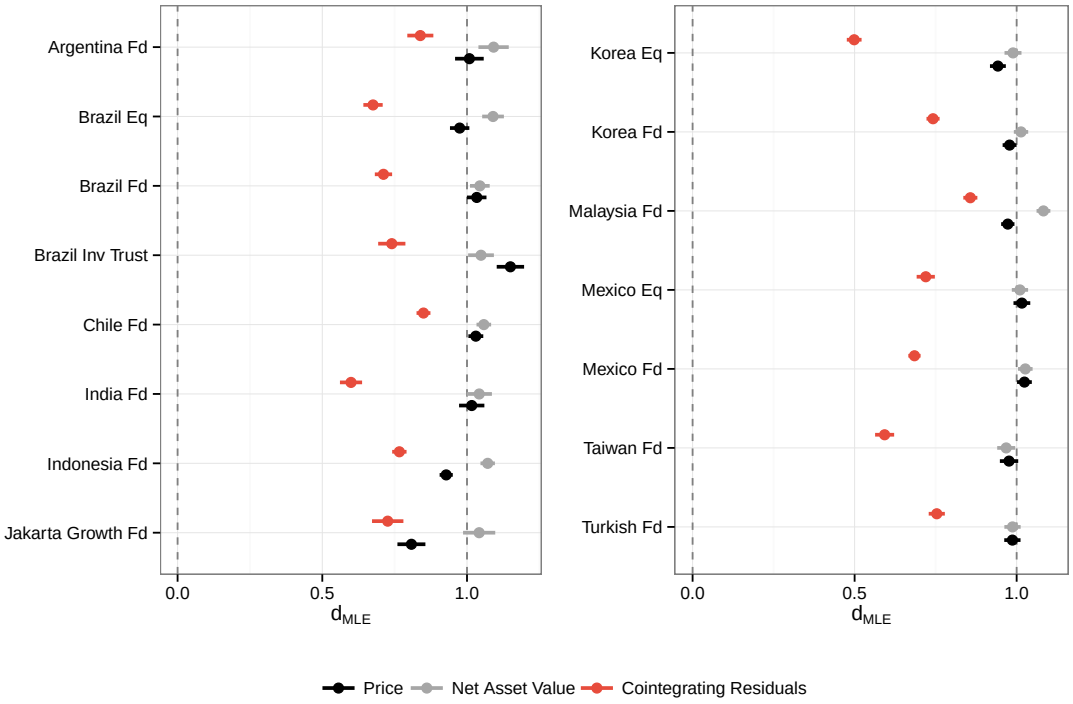

Figure 3. Fractional cointegration of country fund prices and net asset values. Daily country fund data. Maximum likelihood estimates of the fractional differencing parameter, *d*. Horizontal bars are 95% asymptotic confidence intervals.

a few instances, the estimates lie slightly below or above one. More importantly, *d* estimates for the cointegrating residuals can be safely distinguished from both zero and one in every case and are less than the estimated order of integration of the parent series in virtually all cases. The only borderline case is the Jakarta Growth Fund, for which *d* estimates for price and the cointegrating residuals are similar to each other when using daily data. Additionally, results using the GPH estimator for *d*, reported in tables A7 and A8 in the Data Appendix, yield the same conclusions.

There is clear evidence of fractional cointegration in country fund prices and NAVs. Estimates of the fractional integration parameter for the cointegrating residuals range from 0.55 (Brazil Investment Trust) to 0.86 (New South Africa Fund) in the weekly data and from 0.50 (Korea Equity Fund) to 0.86 (Malaysia Fund) in the daily data. The magnitude of the estimates indicates that fund prices and NAVs, therefore, are in long-run equilibrium, and that this equilibrium relation is char-

acterized by mean reversion and long memory. This finding is interesting in itself, since it tells us something important about the nature of interdependence between foreign and domestic investors' market behavior. Because the two series track the value of the same underlying assets—an equity portfolio from a particular emerging market—they largely reflect fundamentals and thus do not drift apart from each other for too long. Any mispricing of a given country's equities on the part of either domestic or international investors is eventually corrected. However, the finding of fractional cointegration also implies that deviations from equilibrium are not corrected immediately; market participants take some time to rebalance their portfolios after a shock.

In addition, estimation results for the fractional error correction model allow us to assess the direction of interdependence between international and domestic investors. The contagion model predicts that short-run movements in NAVs, which contain more information about domestic events, should help predict fund prices, but not the other way around. The model also implies that foreign investors should be the ones to respond to deviations from equilibrium. Tables 2 and 3 show the results of the fractional error corretion analysis using weekly country fund data. The specifications reported in Tables 2 and 3 differ only in the length of the election window used to code the election dummy, E_t.

The results provide strong evidence of contagion from domestic to international investors. Estimates of the long-run adjustment parameter of fund prices, $\hat{\alpha}_1$, are statistically significant for 19 out of 21 funds, indicating that fund prices systematically adjust to deviations from equilibrium. In contrast, estimates of the long-run adjustment parameter of net asset values, $\hat{\alpha}_2$, are statistically significant for only 4 out of 21 country funds. Long-run adjustment of prices to NAVs, therefore, seems to be the rule, while long-run adjustment of NAVs to prices is the exception. Furthermore, in those cases in which both prices and NAVs respond to deviations from equilibrium, the price adjustment parameters are larger than the NAV parameters, suggesting that prices are more sensitive to those deviations.

When it comes to short-run adjustment, the results also lend support to the information asymmetry story. I find evidence of short-run adjustment of prices to NAVs in 9 out 21 funds, while NAVs adjust to prices in only 4 cases. The proportion of cases in which there is any sort of short-run adjustment is much lower than that in which we observe long-run adjustment. This suggests that the latter is a more pervasive mechanism in the interdependent relationship between international and domestic investors, but also that domestic investors often influence the market behavior of foreigners both in the short and the long run.

The weekly analysis in Tables 2 and 3 also yield interesting results for the hypothesis about differential sensitivities to domestic political risk. Regardless of the length of the pre-election window, NAVs show detectable reactions to elections

Table 2. Fractional error correction model of emerging market closed-end country funds (weekly data, 1-month electoral period).

$$\Delta P_t = \phi_1 + \alpha_1[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{1i}\Delta N_{t-i} + \sum_{i=1}^{L}\omega_{1i}\Delta P_{t-i} + \beta_1 E_t + \upsilon_{1t},$$

$$\Delta N_t = \phi_2 + \alpha_2[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{2i}\Delta P_{t-i} + \sum_{i=1}^{L}\omega_{2i}\Delta N_{t-i} + \beta_2 E_t + \upsilon_{2t}$$

				Long-run adjustment				*Short-run adjustment* H_0: $\gamma' = 0$		*Electoral period*			
Country Fund	N	$\hat{d}$	*Lags*	$\hat{\alpha}_1$	t	$\hat{\alpha}_2$	t	F_1	F_2	$\hat{\beta}_1$	t	$\hat{\beta}_2$	t
Argentina Fd	530	0.84	1	**-1.00**	-4.04	0.05	0.29	1.31	0.01	**0.02**	1.79	0.01	1.29
Brazil Eq	682	0.62	1	**-0.50**	-4.90	0.10	1.00	**7.56**	0.72	0.00	0.34	-0.01	-0.57
Brazil Fd	949	0.73	1	**-0.36**	-3.13	0.22	1.91	2.58	1.05	0.00	-0.08	0.00	-0.13
Brazil Inv Trust	273	0.55	1	**-0.57**	-4.57	0.24	1.93	2.58	2.51	0.00	-0.38	-0.01	-1.50
Chile Fd	1304	0.80	1	**-0.58**	-3.86	0.04	0.33	1.46	1.46	0.01	0.73	0.00	-0.23
First Philippine Fd	711	0.74	2	**-1.08**	-5.58	0.21	1.44	**9.38**	**6.62**	0.02	1.37	0.01	0.90
India Fd	1078	0.74	1	**-0.36**	-3.02	0.05	0.49	0.00	0.04	0.01	1.46	0.00	0.59
India Growth Fd	772	0.80	1	-0.20	-1.42	**0.46**	3.63	0.01	**4.55**	0.00	0.07	-0.01	-1.30
Indonesia Fd	1280	0.74	3	**-0.93**	-4.79	-0.06	-0.43	**7.18**	**7.19**	0.01	0.86	0.01	1.40
Jakarta Growth Fd	583	0.69	1	**-0.44**	-3.54	**0.19**	2.15	**4.29**	1.41	0.03	1.11	0.03	1.51
Korea Eq	1089	0.77	1	-0.22	-1.79	0.22	1.75	0.05	3.72	-0.01	-0.89	**-0.04**	-4.48
Korea Fd	1277	0.77	1	**-0.47**	-2.76	0.07	0.45	1.09	0.68	-0.01	-1.22	**-0.04**	-3.16
Malaysia Fd	1285	0.77	1	**-0.28**	-2.61	**0.25**	3.04	1.67	3.44	-0.01	-0.54	0.00	-0.51
Mexico Eq	1210	0.69	1	**-0.70**	-6.10	-0.15	-1.46	**8.03**	**4.40**	0.01	0.81	0.01	0.98
Mexico Fd	1380	0.71	1	**-0.33**	-3.26	**0.18**	2.01	1.74	0.10	0.00	0.42	0.01	1.23
New South Africa Fd	275	0.86	1	**-1.08**	-2.25	-0.62	-1.34	0.96	1.99	**0.02**	2.15	**0.02**	1.69
Taiwan Eq	303	0.79	1	**-0.61**	-2.17	0.10	0.35	0.27	0.00	-0.01	-0.96	-0.01	-0.76
Taiwan Fd	1397	0.83	1	**-0.34**	-2.91	0.11	1.03	**4.28**	0.03	0.00	0.19	0.00	0.47
Thai Capital	1212	0.77	1	**-0.66**	-5.48	0.03	0.42	**10.98**	0.37	-0.01	-0.62	0.00	-0.05
Thai Fd	1390	0.82	1	**-0.44**	-3.05	0.15	1.32	**4.08**	0.66	0.00	0.43	0.01	1.64
Turkish Fd	1295	0.70	1	**-0.48**	-4.83	0.15	1.38	**4.41**	0.01	0.00	0.09	0.00	-0.40

Entries in bold are statistically significant at the 5% level, except for estimates of the election parameter, β, which are reported at the 10% level.

$\hat{d}$ is the Sowell (1992) maximum likelihood estimate of the fractional integration parameter.

Table 3. Fractional error correction model of emerging market closed-end country funds (weekly data, 2-month electoral period).

$$\Delta P_t = \phi_1 + \alpha_1[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{1i}\Delta N_{t-i} + \sum_{i=1}^{L}\omega_{1i}\Delta P_{t-i} + \beta_1 E_t + \upsilon_{1t},$$

$$\Delta N_t = \phi_2 + \alpha_2[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{2i}\Delta P_{t-i} + \sum_{i=1}^{L}\omega_{2i}\Delta N_{t-i} + \beta_2 E_t + \upsilon_{2t}$$

				Long-run adjustment				*Short-run adjustment* H_0: $\gamma' = 0$		*Electoral period*			
Country Fund	N	$\hat{d}$	*Lags*	$\hat{\alpha}_1$	t	$\hat{\alpha}_2$	t	F_1	F_2	$\hat{\beta}_1$	t	$\hat{\beta}_2$	t
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Korea Eq	1089	0.77	1	-0.22	-1.79	0.20	1.56	0.05	3.27	-0.01	-1.39	**-0.03**	-4.16
Korea Fd	1277	0.77	1	**-0.47**	-2.77	0.06	0.36	1.09	0.58	-0.01	-1.63	**-0.03**	-3.12
Malaysia Fd	1285	0.77	1	**-0.29**	-2.64	**0.25**	3.02	1.74	3.37	-0.01	-0.84	0.00	-0.56
Mexico Eq	1210	0.69	1	**-0.70**	-6.10	-0.14	-1.44	**8.01**	**4.31**	0.01	0.91	0.00	0.46
Mexico Fd	1380	0.71	1	**-0.32**	-3.22	**0.18**	2.08	1.63	0.13	0.01	1.15	**0.01**	1.93
New South Africa Fd	275	0.86	1	**-1.21**	-2.47	-0.72	-1.52	1.47	2.52	0.01	1.44	0.01	1.14
Taiwan Eq	303	0.79	1	**-0.66**	-2.33	0.09	0.32	0.45	0.00	0.00	0.36	0.00	-0.45
Taiwan Fd	1397	0.83	1	**-0.35**	-2.93	0.10	1.03	**4.34**	0.03	0.01	0.89	0.00	0.68
Thai Capital	1212	0.77	1	**-0.66**	-5.50	0.04	0.43	**11.10**	0.37	-0.01	-0.97	0.00	0.07
Thai Fd	1390	0.82	1	**-0.45**	-3.09	0.15	1.28	**4.25**	0.59	0.00	-0.33	0.00	1.02
Turkish Fd	1295	0.70	1	**-0.48**	-4.81	0.16	1.43	**4.36**	0.01	0.00	-0.38	-0.01	-1.58

Entries in bold are statistically significant at the 5% level, except for estimates of the election parameter, β, which are reported at the 10% level.

$\hat{d}$ is the Sowell (1992) maximum likelihood estimate of the fractional integration parameter.

more often than fund prices. The difference, however, is not stark. When using a pre-election window of one month, we only see a statistically significant response of prices to elections in two cases; for NAVs, I find a discernible market reaction in three cases. Using a two-month pre-election window, the number of cases in which NAVs respond to elections increases to five, while the number of instances of price responses to elections stays the same. The analysis of weekly data, therefore, provides initial evidence that local investors are more sensitive to domestic political risk than international investors. The number of cases in which we find discernible market reactions to elections, however, is low. This is not unsurprising. First, many elections are relatively predictable, and often large changes in the orientation of economic policy are not at stake. Second, the analysis in Tables 2 and 3 is ill-equipped to detect market reactions to elections, as discussed above.

The FECM analysis using daily country fund data strengthens the results. Tables 4 and 5 report results from the fractional error correction analysis combined with the event study design. A FECM was estimated for each election and country fund in the sample, and pre-election windows of one (Table 4) and two months (Table 5) were used. The results again provide strong support for the prediction of contagion from domestic to international investors. In 27 out of 32 cases, the coefficient estimate for the long-run adjustment of prices to NAVs, $\hat{\alpha}_1$, is statistically significant, thus indicating that fund prices systematically adjust to deviations from equilibrium. For NAVs, the same is true in 13 (Table 4) or 14 (Table 5) out of 32 cases. In those cases in which both prices and NAVs adjust to deviations, the coefficient for the price adjustment tends to be greater than that of the NAV adjustment, again suggesting that international investors are more responsive to deviations from equilibrium than domestic investors.

The pattern of short-term adjustment is also similar to the weekly analysis. I find that fund prices respond to short-run movements in NAVs in 13 (Table 4) and 12 (Table 5) out of 32 cases, while NAVs respond to short-run movements in prices in 8 (Table 4) and 9 (Table 5) out of 32 cases. These results reinforce the interpretation that information flows from domestic to international investors and that long-run adjustment is the dominant mechanism in this interdependence. Figure 4 shows the propagation over time of the effect on country fund prices of a shock to the NAV.

Moreover, the event study provides further evidence that domestic investors are more sensitive to variation in domestic political risk than foreign investors. Table 4 reports estimated cumulative abnormal returns ($\widehat{CAR}$) over a one-month pre-election window. $\widehat{CAR}$ provide an aggregate, intuitive measure of the abnormal behavior of equity markets prior to an election. In the analysis, the net asset value responded to upcoming elections in six cases, while fund prices only displayed any discernible abnormal behavior in two instances. The pattern is un-

Table 4. Fractional error correction model of closed-end country funds with election event study (daily data, 1-month pre-election event window).

$$\Delta P_t = \phi_1 + \alpha_1[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{1i}\Delta N_{t-i} + \sum_{i=1}^{L}\omega_{1i}\Delta P_{t-i} + \sum_{\tau=T_1}^{T_2}\beta_{1,\tau}D_{\tau,t} + \upsilon_{1t},$$

$$\Delta N_t = \phi_2 + \alpha_2[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{2i}\Delta P_{t-i} + \sum_{i=1}^{L}\omega_{2i}\Delta N_{t-i} + \sum_{\tau=T_1}^{T_2}\beta_{2,\tau}D_{\tau,t} + \upsilon_{2t}$$

						Long-run adjustment				*Short-run adjustment* H_0: $\gamma' = 0$		*Event study* H_0: $CAR = 0$ ($CAR \equiv \sum \beta_\tau$)			
Country	*Election*	*Fund*	N	$\hat{d}$	*Lags*	$\hat{\alpha}_1$	t	$\hat{\alpha}_2$	t	F_1	F_2	$\widehat{CAR}_1$	χ^2	$\widehat{CAR}_2$	χ^2
Argentina	1999	Argentina Fd	481	0.66	1	**-0.65**	-3.85	0.00	0.04	2.38	0.23	-0.01	0.01	0.05	0.66
Brazil	2002	Brazil Eq	520	0.68	2	**-0.71**	-3.48	0.04	0.25	**3.29**	1.30	0.04	0.13	-0.04	0.21
Brazil	2002	Brazil Fd	520	0.70	2	**-0.61**	-2.29	0.25	0.88	1.02	1.49	-0.06	0.40	-0.07	0.47
Brazil	2014	Brazil Inv Trust	520	0.73	2	**-0.76**	-4.55	-0.11	-0.61	**4.73**	0.63	-0.03	0.44	**-0.14**	11.29
Chile	2009	Chile Fd	521	0.73	2	0.01	0.05	**0.36**	2.19	**17.81**	**3.32**	0.02	0.05	0.03	0.22
India	2014	India Fd	520	0.60	1	**-0.48**	-4.03	**0.30**	2.37	2.56	1.72	-0.01	0.08	0.02	0.12
Indonesia	1999	Jakarta Growth Fd	336	0.72	1	-0.25	-1.36	**0.31**	2.94	0.37	**5.67**	0.01	0.00	0.10	0.49
Indonesia	2004	Indonesia Fd	521	0.76	1	**-0.81**	-5.26	-0.04	-0.53	**17.07**	0.56	0.18	1.09	0.03	0.12
Indonesia	2009	Indonesia Fd	522	0.75	2	**-1.00**	-4.49	0.20	1.10	**11.13**	**4.14**	-0.03	0.05	0.05	0.25
Indonesia	2014	Indonesia Fd	522	0.61	1	**-0.69**	-4.70	-0.01	-0.09	3.80	0.27	0.03	0.23	0.04	0.37
South Korea	1997	Korea Fd	251	0.83	1	-0.38	-1.12	0.23	0.86	0.62	0.01	0.14	0.71	**-0.58**	18.58
South Korea	2002	Korea Eq	469	0.36	1	**-0.69**	-9.35	**0.22**	3.26	**23.87**	0.56	0.00	0.00	0.01	0.02
South Korea	2002	Korea Fd	522	0.58	1	**-0.73**	-6.26	0.07	0.59	**12.32**	0.18	0.01	0.01	0.01	0.04
South Korea	2007	Korea Eq	522	0.66	1	**-0.42**	-4.34	0.11	1.44	3.73	0.30	-0.03	0.17	-0.02	0.20
South Korea	2007	Korea Fd	522	0.71	1	**-0.79**	-4.15	-0.10	-0.55	**6.91**	0.11	**-0.54**	65.26	**-0.51**	66.55
South Korea	2012	Korea Eq	522	0.32	1	**-0.38**	-7.09	**0.36**	4.73	0.03	**12.22**	-0.02	0.15	-0.01	0.03
South Korea	2012	Korea Fd	522	0.47	1	**-0.39**	-3.46	**0.35**	2.93	0.34	**10.59**	0.01	0.03	0.02	0.13
Malaysia	1995	Malaysia Fd	521	0.87	1	**-0.65**	-2.31	-0.06	-0.30	**4.41**	0.70	-0.02	0.06	-0.03	0.21
Malaysia	1999	Malaysia Fd	520	0.93	1	0.11	0.21	**1.08**	2.78	0.06	**6.93**	-0.10	0.39	-0.07	0.35
Malaysia	2004	Malaysia Fd	521	0.68	1	**-0.57**	-4.76	-0.03	-0.55	3.36	2.60	-0.04	0.16	0.04	1.17
Malaysia	2008	Malaysia Fd	522	0.61	1	**-0.62**	-5.11	0.14	1.89	1.27	1.26	-0.04	0.31	-0.06	2.21
Mexico	2000	Mexico Fd	521	0.63	1	**-0.36**	-2.74	**0.41**	3.12	1.05	**4.78**	0.05	0.33	0.06	0.47
Mexico	2006	Mexico Eq	520	0.80	2	**-1.33**	-3.56	-0.42	-1.25	**7.32**	2.00	-0.02	0.09	-0.03	0.16
Mexico	2006	Mexico Fd	520	0.63	1	-0.25	-1.74	**0.34**	2.18	0.52	0.68	0.05	0.99	-0.02	0.09
Mexico	2012	Mexico Eq	521	0.56	1	**-0.50**	-5.87	**0.32**	3.42	**13.57**	**7.77**	0.02	0.20	**0.08**	3.18
Mexico	2012	Mexico Fd	521	0.61	2	**-0.54**	-3.51	0.25	1.47	**7.21**	1.50	**0.06**	2.79	**0.10**	5.63
Taiwan	2008	Taiwan Fd	522	0.62	1	**-0.30**	-2.90	**0.28**	2.38	1.04	2.55	0.10	2.29	**0.16**	4.51
Taiwan	2012	Taiwan Fd	521	0.48	1	**-0.34**	-5.57	**0.35**	4.96	1.51	**8.45**	0.04	0.91	-0.07	2.12
Turkey	2002	Turkish Fd	435	0.67	1	**-0.28**	-2.44	**0.50**	2.98	0.34	2.71	0.02	0.03	0.02	0.01
Turkey	2007	Turkish Fd	520	0.78	1	**-0.53**	-3.51	0.06	0.34	**7.77**	0.04	0.06	0.38	0.05	0.23
Turkey	2011	Turkish Fd	520	0.66	1	**-0.50**	-2.86	0.18	1.00	1.22	1.08	0.03	0.28	0.03	0.24
Turkey	2014	Turkish Fd	520	0.76	1	**-0.75**	-2.10	-0.32	-0.79	0.68	0.04	0.00	0.00	0.02	0.03

Entries in bold are statistically significant at the 5% level, except for *CAR* estimates, which are shown at the 10% level.

$\hat{d}$ is the Sowell (1992) maximum likelihood estimate of the fractional integration parameter.

Table 5. Fractional error correction model of closed-end country funds with election event study (daily data, 2-month pre-election event window).

$$\Delta P_t = \phi_1 + \alpha_1[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{1i}\Delta N_{t-i} + \sum_{i=1}^{L}\omega_{1i}\Delta P_{t-i} + \sum_{\tau=t_1}^{t_2}\beta_{1,\tau}D_{\tau,t} + \upsilon_{1t},$$

$$\Delta N_t = \phi_2 + \alpha_2[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{2i}\Delta P_{t-i} + \sum_{i=1}^{L}\omega_{2i}\Delta N_{t-i} + \sum_{\tau=t_1}^{t_2}\beta_{2,\tau}D_{\tau,t} + \upsilon_{2t}$$

						Long-run adjustment				*Short-run adjustment* H_0: $\gamma' = 0$		*Event study* H_0: $CAR = 0$ ($CAR \equiv \sum \beta_\tau$)			
Country	*Election*	*Fund*	*N*	$\hat{d}$	*Lags*	$\hat{\alpha}_1$	*t*	$\hat{\alpha}_2$	*t*	F_1	F_2	$\widehat{CAR}_1$	χ^2	$\widehat{CAR}_2$	χ^2
Argentina	1999	Argentina Fd	481	0.66	1	**-0.63**	-3.64	0.01	0.11	1.95	0.32	0.05	0.17	0.12	1.84
Brazil	2002	Brazil Eq	520	0.68	2	**-0.64**	-3.09	0.01	0.07	2.23	2.85	-0.10	0.56	-0.05	0.20
Brazil	2002	Brazil Fd	520	0.70	2	**-0.57**	-2.16	0.27	0.97	2.03	2.63	-0.12	0.99	-0.06	0.21
Brazil	2014	Brazil Inv Trust	520	0.73	2	**-0.90**	-5.03	-0.03	-0.16	**6.72**	0.35	**-0.10**	2.89	-0.08	1.70
Chile	2009	Chile Fd	521	0.73	2	0.01	0.06	**0.36**	2.14	**17.10**	**3.20**	0.04	0.07	0.07	0.52
India	2014	India Fd	520	0.60	1	**-0.47**	-3.85	**0.31**	2.37	2.37	1.71	0.03	0.26	0.03	0.18
Indonesia	1999	Jakarta Growth Fd	336	0.72	1	-0.29	-1.53	**0.34**	2.86	0.85	**5.02**	**0.88**	6.19	0.28	1.58
Indonesia	2004	Indonesia Fd	521	0.76	1	**-0.81**	-5.17	-0.04	-0.53	**16.54**	0.55	0.07	0.08	-0.02	0.03
Indonesia	2009	Indonesia Fd	522	0.75	2	**-1.05**	-4.57	0.17	0.90	**11.80**	**4.02**	0.15	0.77	0.10	0.50
Indonesia	2014	Indonesia Fd	522	0.61	1	**-0.69**	-4.59	-0.02	-0.11	3.48	0.27	0.03	0.08	0.02	0.05
South Korea	1997	Korea Fd	251	0.83	1	0.10	0.27	**1.00**	2.92	0.27	**4.50**	-0.23	1.45	**-1.03**	34.47
South Korea	2002	Korea Eq	469	0.36	1	**-0.67**	-8.81	**0.24**	3.40	**20.79**	0.74	-0.05	0.14	-0.12	0.93
South Korea	2002	Korea Fd	522	0.58	1	**-0.70**	-5.84	0.10	0.88	**10.14**	0.42	-0.01	0.00	-0.07	0.52
South Korea	2007	Korea Eq	522	0.66	1	**-0.43**	-4.31	0.12	1.64	2.69	0.79	0.06	0.43	-0.07	1.14
South Korea	2007	Korea Fd	522	0.71	1	**-0.81**	-4.09	-0.10	-0.53	**6.23**	0.04	**-0.46**	24.24	**-0.50**	31.54
South Korea	2012	Korea Eq	522	0.32	1	**-0.39**	-7.10	**0.36**	4.65	0.09	**11.90**	-0.02	0.17	-0.06	0.60
South Korea	2012	Korea Fd	522	0.47	1	**-0.41**	-3.51	**0.36**	2.87	0.26	**10.29**	0.00	0.01	0.06	0.63
Malaysia	1995	Malaysia Fd	521	0.87	1	**-0.66**	-2.27	-0.12	-0.56	**4.30**	1.31	-0.10	0.52	-0.04	0.15
Malaysia	1999	Malaysia Fd	520	0.93	1	0.08	0.15	**1.11**	2.77	0.04	**6.88**	-0.17	0.52	0.04	0.04
Malaysia	2004	Malaysia Fd	521	0.68	1	**-0.51**	-4.25	-0.01	-0.22	1.85	2.27	-0.07	0.28	0.07	1.45
Malaysia	2008	Malaysia Fd	522	0.61	1	**-0.53**	-4.22	**0.16**	2.16	0.79	1.68	0.05	0.30	-0.06	0.93
Mexico	2000	Mexico Fd	521	0.63	1	**-0.40**	-2.93	**0.40**	2.94	1.43	3.77	-0.01	0.01	0.08	0.44
Mexico	2006	Mexico Eq	520	0.80	2	**-1.24**	-3.26	-0.46	-1.35	**6.11**	2.50	-0.01	0.02	-0.03	0.07
Mexico	2006	Mexico Fd	520	0.63	1	-0.27	-1.74	0.13	0.80	0.36	0.02	0.06	0.83	0.00	0.00
Mexico	2012	Mexico Eq	521	0.56	1	**-0.45**	-5.19	**0.35**	3.58	**8.27**	**8.30**	0.04	0.57	0.05	0.61
Mexico	2012	Mexico Fd	521	0.61	2	**-0.55**	-3.47	0.25	1.47	**6.88**	1.50	0.05	0.81	**0.11**	3.33
Taiwan	2008	Taiwan Fd	522	0.62	1	**-0.30**	-2.84	**0.29**	2.44	1.22	2.86	0.10	1.18	0.15	1.99
Taiwan	2012	Taiwan Fd	521	0.48	1	**-0.34**	-5.44	**0.35**	4.94	0.80	**10.51**	-0.03	0.39	**-0.13**	4.21
Turkey	2002	Turkish Fd	435	0.67	1	**-0.27**	-2.26	**0.51**	2.92	0.22	2.44	0.06	0.10	0.23	0.65
Turkey	2007	Turkish Fd	520	0.78	1	**-0.52**	-3.39	0.06	0.34	**7.16**	0.02	0.00	0.00	0.03	0.04
Turkey	2011	Turkish Fd	520	0.66	1	**-0.49**	-2.80	0.18	0.98	1.07	1.05	-0.02	0.05	-0.01	0.01
Turkey	2014	Turkish Fd	520	0.76	1	**-0.75**	-2.04	-0.33	-0.82	0.55	0.04	-0.05	0.16	-0.03	0.04

Entries in bold are statistically significant at the 5% level, except for *CAR* estimates, which are shown at the 10% level.

$\hat{d}$ is the Sowell (1992) maximum likelihood estimate of the fractional integration parameter.

changed when using a two-month pre-election window, as shown in Table 5. We can discern NAV reactions in four cases and price reactions in two. The lower number of detected market reactions when using a two-month window is unsurprising given that the statistical power of an event study is decreasing in the length of the event window (Campbell, Lo, and MacKinlay, 1997).

The final set of results corroborates the pattern of dynamic interdependence between domestic and international investors while providing a complementary view of how these investors acquire and process information on the domestic politics of emerging economies. The analysis reported in Table 6 takes advantage of the availability of daily country fund data and, on top of estimating contagion dynamics between domestic and foreign investors, also assesses how these two groups respond to the release of new polling information in years of presidential election. The inclusion of a series of entropy measurements throughout the election year tests the extent to which foreign and domestic investors closely track short-run developments in the domestic politics of emerging markets.

With respect to the contagion dynamics, I find that fund prices respond to deviations from equilibrium in election years in 8 out of 14 cases, while NAVs only show long-run adjustment in 2 cases. When it comes to short-run adjustment, the pattern repeats itself: fund prices respond to short-run movements in NAVs in 4 cases, while NAVs respond to short-run movements in prices in none of the cases. The lower proportion of statistically significant results, when compared to the previous analyses, seems to be a function of smaller sample sizes when the analysis is restricted to election years. Nonetheless, these results further confirm the finding of contagion from domestic to international financial investors.

Results for the entropy measure are also in line with the previous results. We can statistically discern NAV responses to short-run changes in electoral uncertainty in 5 of the 14 cases, but no reaction is detected for country fund prices. Interestingly, all five cases of significant domestic investor reaction involve Brazilian presidential elections. Among included emerging markets, Brazil has possibly the highest level of economic policy volatility (Fatás and Mihov, 2013). Moreover, the country has the deepest financial market in the sample. BM&F BOVESPA, the São Paulo Stock Exchange and Futures Market, is among the main exchanges in the emerging world, ranking only behind the Shanghai Stock Exchange, the Bombay Stock Exchange, and the National Stock Exchange of India in terms of market capitalization.[7] Large and liquid financial markets make it easier for investors to rebalance their portfolios (Eichengreen and Gupta, Forthcoming). Therefore, the combination of high economic policy volatility with a deep equity market creates a strong candidate for high financial market sensitivity to electoral uncertainty, and might help explain the dominance of Brazilian elections among the significant re-

[7] Monthly Reports, World Federation of Exchanges. May 2014.

Table 6. Fractional error correction model of emerging market closed-end country funds (daily data).

$$\Delta P_t = \phi_1 + \alpha_1[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{1i}\Delta N_{t-i} + \sum_{i=1}^{L}\omega_{1i}\Delta P_{t-i} + \beta_1 \Delta Entropy_t + \upsilon_{1t},$$

$$\Delta N_t = \phi_2 + \alpha_2[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{2i}\Delta P_{t-i} + \sum_{i=1}^{L}\omega_{2i}\Delta N_{t-i} + \beta_2 \Delta Entropy_t + \upsilon_{2t}$$

						Long-run adjustment				*Short-run adjustment* H_0: $\gamma' = 0$		*Entropy*			
Country	*Election*	*Fund*	N	$\hat{d}$	*Lags*	$\hat{\alpha}_1$	t	$\hat{\alpha}_2$	t	F_1	F_2	$\hat{\beta}_1$	t	$\hat{\beta}_2$	t
Brazil	1998	Brazil Eq	145	0.67	1	-0.47	-1.46	0.23	1.47	2.13	2.11	0.14	0.40	**0.32**	1.90
Brazil	1998	Brazil Fd	147	0.61	1	-0.72	-1.80	0.17	0.63	**4.73**	0.24	-0.04	-0.14	**0.33**	1.86
Brazil	2002	Brazil Eq	148	0.49	1	**-0.69**	-4.11	0.14	0.79	1.76	0.08	-0.02	-0.34	**0.18**	2.43
Brazil	2002	Brazil Fd	148	0.35	1	**-0.53**	-2.42	0.31	1.18	0.20	1.99	0.11	1.65	**0.19**	2.32
Brazil	2010	Brazil Inv Trust	134	0.93	1	-0.39	-0.46	-0.47	-0.59	0.13	0.31	0.03	0.54	**0.10**	2.02
Brazil	2014	Brazil Inv Trust	179	0.68	1	-0.32	-1.65	-0.09	-0.37	0.01	0.01	-0.04	-0.73	0.03	0.46
Chile	2009	Chile Fd	271	0.68	1	**-0.67**	-3.72	0.02	0.18	1.80	0.39	0.02	0.08	0.12	0.89
South Korea	2012	Korea Eq	236	0.42	1	**-0.33**	-3.30	**0.24**	2.12	0.83	0.70	-0.01	-0.25	0.03	0.80
South Korea	2012	Korea Fd	236	0.53	1	**-0.69**	-4.97	0.03	0.25	**12.49**	0.75	0.00	0.15	0.02	0.50
Mexico	2006	Mexico Eq	169	0.82	1	-1.21	-1.75	-0.74	-1.21	2.43	1.19	-0.32	-0.64	0.46	1.02
Mexico	2006	Mexico Fd	169	0.51	1	-0.22	-0.99	**0.59**	2.56	0.26	1.23	0.26	0.86	0.15	0.49
Mexico	2012	Mexico Eq	260	0.57	1	**-0.79**	-5.27	0.01	0.04	**16.49**	0.11	0.00	-0.07	-0.02	-0.70
Mexico	2012	Mexico Fd	260	0.48	1	**-0.68**	-5.46	0.19	1.27	**14.35**	0.08	0.00	0.15	-0.01	-0.40
Turkey	2011	Turkish Fd	259	0.54	1	**-0.53**	-2.03	0.19	0.71	0.07	0.42	0.02	0.10	0.06	0.29

Entries in bold are statistically significant at the 5% level, except for β estimates, which are shown at the 10% level.

$\hat{d}$ is the Sowell (1992) maximum likelihood estimate of the fractional integration parameter.

sults.

Taken together, these results paint a clear picture of the dynamics of global financial market reactions to country-level politics. The disaggregation of financial market behavior between domestic and international investors allows us to observe contagion dynamics between these two groups and sheds light on an important mechanism of information transmission in globally-integrated capital markets. Domestic investors appear as major drivers of market reactions to politics in emerging economies, as their behavior anticipates that of international investors in both the short and the long run. This is consistent with a model of information asymmetry in which local investors have lower information costs and are on average better informed than foreign investors. The results show that domestic investors immediately respond to contemporaneous shocks in electoral uncertainty and transmit these shocks to international investors with a delay, as the latter observe domestic market activity and then adjust their portfolios accordingly. International investors do not seem to respond to contemporaneous shocks; rather, they follow the lead of better informed domestic investors. Moreover, having higher exposure to domestic risks, local investors tend to respond with a greater rebalancing of their portfolios, which is reflected in greater abnormal performance during elections relative to international investors.

CONCLUSION

Political economists have accumulated considerable knowledge about how global markets price politics, but little attention has been paid to the dynamic mechanisms behind the start and spread of global market responses to politics. I have offered an account based on information asymmetries between domestic and international investors that predicts that shocks to country-level political risk in emerging markets are transmitted from the former to the latter. This contagion process predicts that better-informed domestic investors closely follow political risk fundamentals, and less well-informed international investors extract information from the market behavior of local investors. In other words, global capital markets act as an information transmission mechanism from local to foreign investors.

The analysis took advantage of closed-end country fund data, which afford a unique opportunity to disaggregate the behavior of domestic and foreign investors. Moreover, the availability of high frequency data allowed for the precise identification of market dynamics. Results of the fractional error correction analysis show both short- and long-run contagion from domestic to foreign investors. Interestingly, the analysis of fractional cointegration reveals that, while international investors eventually correct any mispricings of country-level risks, that correction

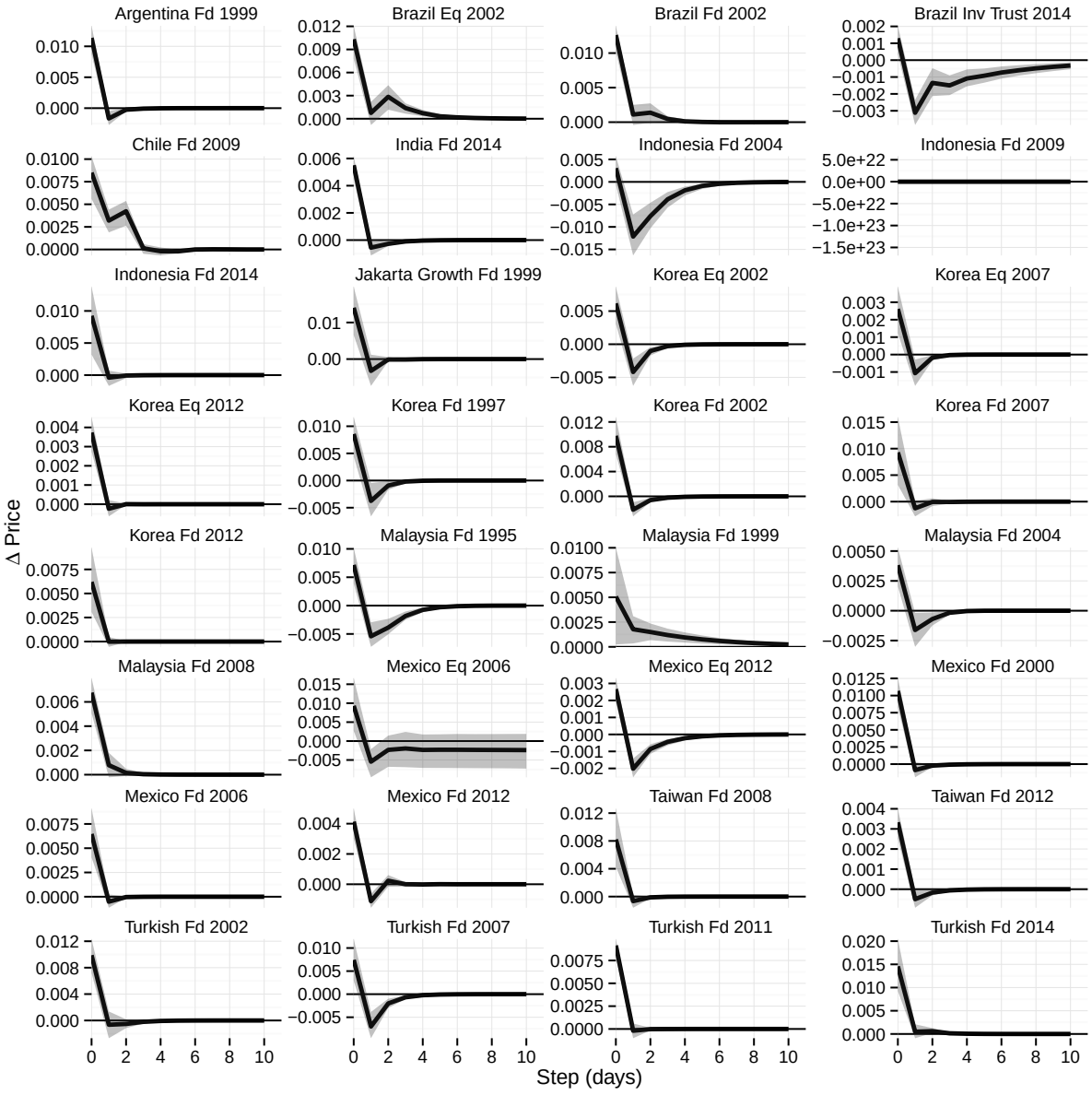

Figure 4. Orthogonalized impulse response functions. Effect on country fund prices of a shock to the NAV. Estimates from the fractional error correction model with election event study (1-month event window) reported in Table 4.

can take quite some time to occur and discrepancies in the reactions of domestic and foreign investors can be persistent.

Are these results generalizable to other financial market segments? The analysis in this article presents direct evidence of domestic-to-foreign contagion in equity markets. Can we expect its conclusions to hold in sovereign debt and currency markets? Data availability is a major constraint, since either price or transaction data disaggregated by domicile are rarely available for any securities. However, there is no reason to assume that the same dynamics would not hold in different security markets. The theoretical model of market contagion is not specific to equity markets, and the same type of information asymmetries should be observed among sovereign debt and currency traders, for example, where information costs are at least as pronounced as in international equity markets. The findings from equity markets, therefore, should provide good hints of what to expect from global market dynamics in markets for different securities.

The results have both political and policy implications. Foreign capital has often been named as one of the culprits in debates of the political and economic costs of capital mobility. Nonetheless, the findings in this study suggest a more nuanced view of the role of foreign capital in constraining the policy autonomy of emerging-market governments. They suggest that an increased presence of foreign investors in emerging markets might not directly increase the probability of crises or speculative attacks initiated by political events. Instead, the analysis indicates that the contribution of international investors might be rather indirect—they amplify the responses of domestic capital markets to politics through a mechanism of contagion.

Moreover, if foreign and domestic investors operate with different information, use different strategies, and respond in different ways to politics in emerging economies, then it matters which audiences governments choose to target when seeking to appease markets. The contagion dynamics analyzed in this paper suggest that domestic investors are a critical audience for governments facing credibility crises and play a central part in shaping global market perceptions of country risk. The findings, therefore, reveal an important new dimension of the confidence game played between governments and markets in the emerging world.

DATA APPENDIX

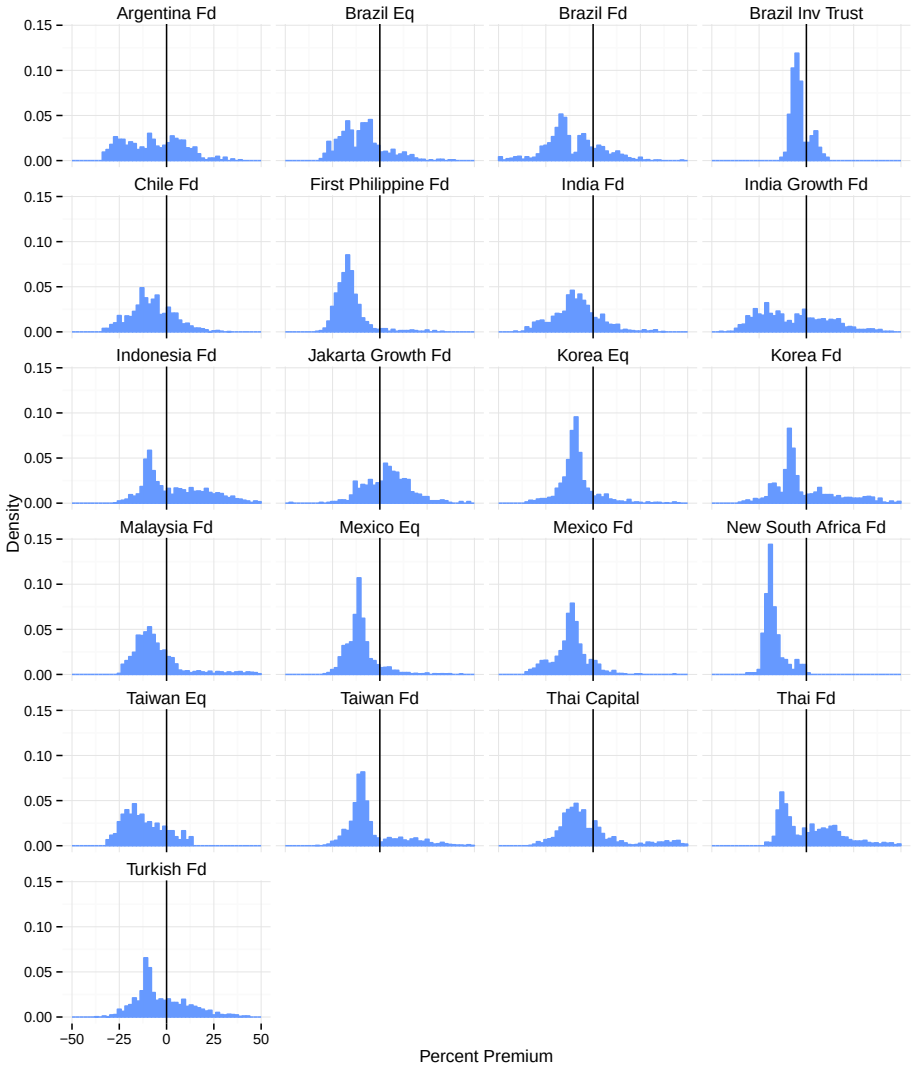

Figure A1. Distribution of closed-end country fund premia.

Table A1. Dickey-Fuller and Augmented Dickey-Fuller Test for Unit Root (Weekly Country Fund Data).

	1% critical value = −3.43											
	Price				*Net Asset Value*				*Cointegrating Residuals*			
	$k=0$	$k=1$	$k=2$	$k=3$	$k=0$	$k=1$	$k=2$	$k=3$	$k=0$	$k=1$	$k=2$	$k=3$
Argentina Fd	-3.00	-3.06	-2.94	-3.07	-1.95	-2.22	-2.50	-2.40	-3.12	-2.97	-2.60	-2.60
Brazil Eq	-1.86	-1.75	-1.80	-1.82	-1.76	-1.71	-1.73	-1.80	-5.66	-4.09	-3.61	-3.33
Brazil Fd	0.81	1.15	0.76	0.78	0.17	0.15	-0.03	-0.13	-4.65	-3.48	-3.14	-3.08
Brazil Inv Trust	-0.82	-0.88	-0.91	-0.62	-0.73	-0.71	-0.78	-0.69	-5.53	-4.40	-3.47	-2.97
Chile Fd	-2.42	-2.25	-2.31	-2.48	-1.98	-2.08	-2.24	-2.44	-5.06	-4.12	-3.93	-3.70
First Philippine Fd	-1.01	-0.97	-1.14	-0.98	-0.25	-0.42	-0.62	-0.64	-6.52	-5.88	-6.92	-5.52
India Fd	-1.72	-1.51	-1.65	-1.77	-1.40	-1.52	-1.74	-1.79	-4.75	-3.43	-3.15	-3.11
India Growth Fd	-2.18	-2.26	-2.41	-2.56	-2.33	-2.28	-2.58	-2.93	-3.74	-3.00	-2.77	-2.71
Indonesia Fd	-2.69	-2.48	-2.47	-2.52	-1.72	-1.71	-1.84	-2.02	-5.55	-4.61	-4.11	-3.69
Jakarta Growth Fd	-1.42	-1.18	-1.31	-1.80	-0.62	-0.64	-0.91	-1.13	-5.45	-4.03	-3.32	-3.44
Korea Eq	-1.66	-1.61	-1.59	-1.68	-1.40	-1.36	-1.41	-1.39	-5.21	-4.15	-3.83	-3.92
Korea Fd	-1.47	-1.48	-1.54	-1.67	-1.30	-1.36	-1.47	-1.63	-3.75	-3.32	-3.11	-3.08
Malaysia Fd	-1.90	-1.73	-1.79	-1.86	-1.32	-1.40	-1.47	-1.51	-5.10	-3.98	-3.60	-3.61
Mexico Eq	-2.13	-1.97	-2.02	-2.10	-1.68	-1.80	-1.99	-2.06	-6.81	-4.85	-4.18	-4.04
Mexico Fd	-2.47	-2.42	-2.49	-2.53	-2.43	-2.47	-2.64	-2.57	-6.05	-4.43	-4.09	-3.83
New South Africa Fd	-1.70	-2.09	-2.35	-2.49	-0.85	-1.33	-1.64	-1.80	-4.68	-3.84	-3.80	-3.93
Taiwan Eq	-0.40	-0.53	-0.81	-0.83	-0.77	-0.94	-0.97	-1.40	-2.89	-2.19	-1.92	-1.73
Taiwan Fd	-3.05	-3.05	-2.79	-2.62	-2.86	-2.84	-2.94	-3.11	-6.31	-5.70	-5.34	-4.91
Thai Capital	-1.80	-1.69	-1.68	-1.69	-1.27	-1.32	-1.36	-1.34	-4.62	-3.73	-3.48	-3.15
Thai Fd	-2.17	-2.01	-2.03	-1.99	-1.49	-1.48	-1.55	-1.51	-4.75	-4.00	-3.84	-3.64
Turkish Fd	-2.43	-2.41	-2.59	-2.62	-2.45	-2.49	-2.74	-2.79	-6.41	-4.80	-4.16	-3.95

Table cells show the *DF* test statistic. H_0: Series is $I(1)$ (random walk with drift). k is the lag length.

Table A2. KPSS Test for Stationarity (Weekly Country Fund Data).

	Price				*Net Asset Value*				*Cointegrating Residuals*			
	$k = 8$		$k = 25$		$k = 8$		$k = 25$		$k = 8$		$k = 25$	
	KPSS	*p*	*KPSS*	*p*	*KPSS*	*p*	*KPSS*	*p*	*KPSS*	*p*	*KPSS*	*p*
Argentina Fd	0.26	0.01	0.11	0.10	0.46	0.01	0.19	0.02	0.28	0.01	0.12	0.09
Brazil Eq	0.74	0.01	0.29	0.01	0.64	0.01	0.26	0.01	1.38	0.01	0.55	0.01
Brazil Fd	0.91	0.01	0.34	0.01	0.71	0.01	0.27	0.01	0.93	0.01	0.34	0.01
Brazil Inv Trust	0.14	0.06	0.06	0.10	0.33	0.01	0.14	0.06	0.96	0.01	0.39	0.01
Chile Fd	0.92	0.01	0.34	0.01	0.83	0.01	0.30	0.01	1.63	0.01	0.62	0.01
First Philippine Fd	1.90	0.01	0.68	0.01	1.93	0.01	0.69	0.01	0.22	0.01	0.10	0.10
India Fd	1.35	0.01	0.47	0.01	1.09	0.01	0.38	0.01	0.75	0.01	0.27	0.01
India Growth Fd	1.20	0.01	0.46	0.01	0.55	0.01	0.21	0.01	0.97	0.01	0.37	0.01
Indonesia Fd	2.45	0.01	0.89	0.01	2.62	0.01	0.94	0.01	0.49	0.01	0.19	0.02
Jakarta Growth Fd	1.59	0.01	0.57	0.01	1.41	0.01	0.49	0.01	0.50	0.01	0.20	0.02
Korea Eq	1.83	0.01	0.64	0.01	1.74	0.01	0.61	0.01	0.78	0.01	0.29	0.01
Korea Fd	1.29	0.01	0.46	0.01	1.59	0.01	0.57	0.01	1.97	0.01	0.74	0.01
Malaysia Fd	1.27	0.01	0.46	0.01	1.29	0.01	0.46	0.01	0.46	0.01	0.18	0.03
Mexico Eq	0.69	0.01	0.26	0.01	0.68	0.01	0.25	0.01	0.41	0.01	0.17	0.03
Mexico Fd	0.58	0.01	0.21	0.01	0.46	0.01	0.16	0.04	1.00	0.01	0.36	0.01
New South Africa Fd	0.72	0.01	0.28	0.01	1.00	0.01	0.38	0.01	0.25	0.01	0.12	0.10
Taiwan Eq	0.49	0.01	0.22	0.01	0.33	0.01	0.14	0.06	0.71	0.01	0.31	0.01
Taiwan Fd	1.78	0.01	0.66	0.01	0.83	0.01	0.31	0.01	1.07	0.01	0.48	0.01
Thai Capital	1.58	0.01	0.59	0.01	1.66	0.01	0.61	0.01	0.70	0.01	0.27	0.01
Thai Fd	1.86	0.01	0.65	0.01	1.82	0.01	0.63	0.01	0.34	0.01	0.13	0.09
Turkish Fd	0.53	0.01	0.20	0.01	0.42	0.01	0.16	0.03	0.62	0.01	0.24	0.01

H_0: Series is $I(0)$ (stationary). k is the lag length.

Table A3. Lo-MacKinlay Variance-Ratio Test for Random Walk (Weekly Country Fund Data).

	Price						*Net Asset Value*						*Cointegrating Residuals*					
	q=2		*q*=4		*q*=8		*q*=2		*q*=4		*q*=8		*q*=2		*q*=4		*q*=8	
	M_2	*p*	M_2	*p*	M_2	*p*	M_2	*p*	M_2	*p*	M_2	*p*	M_2	*p*	M_2	*p*	M_2	*p*
Argentina Fd	13.09	0.00	20.50	0.00	28.88	0.00	14.64	0.00	23.06	0.00	32.78	0.00	14.56	0.00	23.02	0.00	33.18	0.00
Brazil Eq	19.41	0.00	31.01	0.00	45.50	0.00	18.24	0.00	29.19	0.00	42.89	0.00	15.49	0.00	24.47	0.00	35.43	0.00
Brazil Fd	17.32	0.00	27.48	0.00	39.65	0.00	16.88	0.00	26.77	0.00	38.60	0.00	15.43	0.00	24.60	0.00	36.00	0.00
Brazil Inv Trust	10.61	0.00	16.72	0.00	23.87	0.00	9.94	0.00	15.61	0.00	22.16	0.00	8.95	0.00	13.40	0.00	18.34	0.00
Chile Fd	24.76	0.00	39.51	0.00	57.80	0.00	22.54	0.00	35.89	0.00	52.25	0.00	21.10	0.00	33.57	0.00	48.99	0.00
First Philippine Fd	19.01	0.00	30.32	0.00	44.28	0.00	18.60	0.00	29.69	0.00	43.40	0.00	7.83	0.00	11.84	0.00	16.18	0.00
India Fd	24.31	0.00	38.92	0.00	57.28	0.00	23.35	0.00	37.37	0.00	54.92	0.00	17.70	0.00	28.17	0.00	41.00	0.00
India Growth Fd	17.98	0.00	28.68	0.00	41.85	0.00	16.25	0.00	25.86	0.00	37.66	0.00	18.96	0.00	30.23	0.00	44.17	0.00
Indonesia Fd	20.54	0.00	32.86	0.00	48.27	0.00	22.86	0.00	36.60	0.00	53.81	0.00	18.33	0.00	29.09	0.00	42.63	0.00
Jakarta Growth Fd	14.27	0.00	22.65	0.00	32.67	0.00	16.07	0.00	25.54	0.00	36.98	0.00	10.56	0.00	16.69	0.00	23.94	0.00
Korea Eq	23.95	0.00	38.32	0.00	56.31	0.00	21.79	0.00	34.89	0.00	51.35	0.00	12.22	0.00	19.40	0.00	28.05	0.00
Korea Fd	23.86	0.00	38.16	0.00	56.04	0.00	24.45	0.00	39.13	0.00	57.53	0.00	20.37	0.00	32.46	0.00	47.37	0.00
Malaysia Fd	25.12	0.00	40.23	0.00	59.25	0.00	24.08	0.00	38.60	0.00	56.91	0.00	15.84	0.00	25.36	0.00	37.45	0.00
Mexico Eq	19.73	0.00	31.55	0.00	46.33	0.00	20.26	0.00	32.40	0.00	47.51	0.00	11.53	0.00	18.36	0.00	26.67	0.00
Mexico Fd	18.54	0.00	29.48	0.00	42.83	0.00	19.65	0.00	31.26	0.00	45.46	0.00	18.32	0.00	29.15	0.00	42.66	0.00
New South Africa Fd	6.54	0.00	10.36	0.00	15.00	0.00	8.49	0.00	13.37	0.00	18.95	0.00	5.14	0.00	7.75	0.00	10.25	0.00
Taiwan Eq	8.59	0.00	13.24	0.00	17.96	0.00	10.46	0.00	16.43	0.00	23.24	0.00	10.95	0.00	17.08	0.00	24.09	0.00
Taiwan Fd	22.20	0.00	35.47	0.00	51.97	0.00	20.71	0.00	33.12	0.00	48.60	0.00	13.27	0.00	20.87	0.00	29.71	0.00
Thai Capital	21.22	0.00	33.95	0.00	49.90	0.00	22.95	0.00	36.74	0.00	54.07	0.00	18.10	0.00	28.85	0.00	42.36	0.00
Thai Fd	23.41	0.00	37.48	0.00	55.16	0.00	25.23	0.00	40.43	0.00	59.62	0.00	19.66	0.00	31.38	0.00	46.07	0.00
Turkish Fd	26.45	0.00	42.28	0.00	62.01	0.00	27.93	0.00	44.68	0.00	65.57	0.00	19.18	0.00	30.51	0.00	44.44	0.00

H_0: Series is $I(1)$ (random walk with drift). q is the span of differencing.

M_2 is a standardized test statistic with standard normal distribution and is robust to conditional heteroskedasticity.

Table A4. Dickey-Fuller and Augmented Dickey-Fuller Test for Unit Root (Daily Country Fund Data).

	1% critical value = −3.43											
	Price				*Net Asset Value*				*Cointegrating Residuals*			
	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 0$	$k = 1$	$k = 2$	$k = 3$	$k = 0$	$k = 1$	$k = 2$	$k = 3$
Argentina Fd	-2.31	-2.39	-2.27	-2.35	-0.77	-1.14	-1.09	-1.16	-3.33	-2.83	-2.51	-2.40
Brazil Eq	-1.22	-1.14	-1.14	-1.16	-1.46	-1.71	-1.68	-1.73	-4.74	-3.55	-2.88	-2.33
Brazil Fd	1.68	1.51	1.49	1.49	1.18	0.96	0.98	0.97	-4.96	-3.83	-3.31	-3.03
Brazil Inv Trust	-0.57	-0.82	-0.83	-0.90	-0.61	-0.75	-0.84	-0.84	-7.14	-6.17	-5.77	-5.56
Chile Fd	-1.98	-2.18	-2.21	-2.10	-1.60	-1.79	-1.84	-1.76	-4.05	-3.57	-3.27	-3.18
India Fd	-1.41	-1.44	-1.47	-1.46	-1.33	-1.39	-1.41	-1.42	-8.66	-6.15	-5.17	-4.48
Indonesia Fd	-1.63	-1.62	-1.63	-1.60	-1.26	-1.29	-1.31	-1.32	-7.04	-6.10	-5.44	-5.11
Jakarta Growth Fd	-2.06	-1.76	-1.58	-1.55	-1.04	-1.23	-1.07	-1.15	-4.48	-3.79	-3.33	-3.28
Korea Eq	-1.84	-1.78	-1.78	-1.80	-1.77	-1.76	-1.74	-1.74	-16.99	-11.74	-9.15	-7.81
Korea Fd	-1.20	-1.16	-1.14	-1.12	-1.10	-1.10	-1.08	-1.06	-5.86	-4.69	-4.04	-3.79
Malaysia Fd	-1.48	-1.55	-1.45	-1.42	-1.04	-1.17	-1.16	-1.18	-4.84	-4.31	-4.18	-3.86
Mexico Eq	-1.63	-1.72	-1.67	-1.64	-1.41	-1.44	-1.44	-1.46	-10.36	-9.01	-7.85	-6.77
Mexico Fd	-1.93	-1.98	-1.96	-1.94	-1.99	-2.08	-2.08	-2.05	-5.53	-4.10	-3.49	-3.10
Taiwan Fd	-2.28	-2.22	-2.28	-2.28	-2.06	-2.05	-1.99	-2.07	-14.55	-10.74	-9.48	-8.50
Turkish Fd	-2.09	-2.06	-2.07	-2.04	-2.03	-1.98	-2.00	-1.99	-6.99	-5.48	-4.90	-4.59

Table cells show the *DF* test statistic. H_0: Series is $I(1)$ (random walk with drift). k is the lag length.

Table A5. KPSS Test for Stationarity (Daily Country Fund Data).

	Price				*Net Asset Value*				*Cointegrating Residuals*			
	$k = 8$		$k = 25$		$k = 8$		$k = 25$		$k = 8$		$k = 25$	
	KPSS	*p*	*KPSS*	*p*	*KPSS*	*p*	*KPSS*	*p*	*KPSS*	*p*	*KPSS*	*p*
Argentina Fd	0.70	0.01	0.25	0.01	1.63	0.01	0.58	0.01	1.52	0.01	0.54	0.01
Brazil Eq	2.17	0.01	0.74	0.01	1.76	0.01	0.59	0.01	2.47	0.01	0.87	0.01
Brazil Fd	3.74	0.01	1.25	0.01	3.62	0.01	1.21	0.01	0.76	0.01	0.26	0.01
Brazil Inv Trust	0.31	0.01	0.11	0.10	0.70	0.01	0.25	0.01	1.92	0.01	0.75	0.01
Chile Fd	2.41	0.01	0.81	0.01	2.33	0.01	0.78	0.01	1.63	0.01	0.57	0.01
India Fd	1.89	0.01	0.68	0.01	1.87	0.01	0.68	0.01	0.56	0.01	0.23	0.01
Indonesia Fd	1.48	0.01	0.50	0.01	1.41	0.01	0.47	0.01	1.49	0.01	0.52	0.01
Jakarta Growth Fd	1.42	0.01	0.49	0.01	2.22	0.01	0.76	0.01	1.53	0.01	0.58	0.01
Korea Eq	1.90	0.01	0.64	0.01	1.72	0.01	0.58	0.01	2.00	0.01	0.74	0.01
Korea Fd	4.58	0.01	1.51	0.01	4.66	0.01	1.54	0.01	1.87	0.01	0.64	0.01
Malaysia Fd	5.25	0.01	1.70	0.01	5.25	0.01	1.70	0.01	1.26	0.01	0.42	0.01
Mexico Eq	2.59	0.01	0.88	0.01	2.76	0.01	0.93	0.01	0.82	0.01	0.32	0.01
Mexico Fd	1.48	0.01	0.50	0.01	1.66	0.01	0.56	0.01	1.27	0.01	0.44	0.01
Taiwan Fd	1.15	0.01	0.38	0.01	1.13	0.01	0.37	0.01	0.63	0.01	0.25	0.01
Turkish Fd	2.75	0.01	0.92	0.01	2.34	0.01	0.79	0.01	1.84	0.01	0.68	0.01

H_0: Series is $I(0)$ (stationary). k is the lag length.

Table A6. Lo-MacKinlay Variance-Ratio Test for Random Walk (Daily Country Fund Data).

	Price						*Net Asset Value*						*Cointegrating Residuals*					
	q=2		*q=4*		*q=8*		*q=2*		*q=4*		*q=8*		*q=2*		*q=4*		*q=8*	
	M_2	p	M_2	p	M_2	p	M_2	p	M_2	p	M_2	p	M_2	p	M_2	p	M_2	p
Argentina Fd	18.74	0.00	29.84	0.00	43.37	0.00	15.57	0.00	24.74	0.00	35.88	0.00	21.44	0.00	34.20	0.00	50.01	0.00
Brazil Eq	25.71	0.00	41.01	0.00	59.91	0.00	27.66	0.00	44.18	0.00	64.72	0.00	29.46	0.00	46.99	0.00	68.88	0.00
Brazil Fd	25.90	0.00	41.42	0.00	60.80	0.00	26.51	0.00	42.40	0.00	62.26	0.00	28.87	0.00	46.15	0.00	67.87	0.00
Brazil Inv Trust	23.91	0.00	38.21	0.00	56.00	0.00	22.59	0.00	36.06	0.00	52.80	0.00	20.64	0.00	32.30	0.00	46.08	0.00
Chile Fd	44.44	0.00	71.18	0.00	104.84	0.00	39.84	0.00	63.80	0.00	93.92	0.00	34.34	0.00	54.93	0.00	80.78	0.00
India Fd	25.12	0.00	40.25	0.00	59.30	0.00	24.44	0.00	39.15	0.00	57.66	0.00	20.04	0.00	31.68	0.00	45.64	0.00
Indonesia Fd	43.53	0.00	69.77	0.00	102.91	0.00	46.68	0.00	74.82	0.00	110.30	0.00	34.35	0.00	54.82	0.00	80.43	0.00
Jakarta Growth Fd	21.68	0.00	34.58	0.00	50.55	0.00	19.46	0.00	31.05	0.00	45.42	0.00	14.99	0.00	23.77	0.00	34.47	0.00
Korea Eq	40.55	0.00	64.94	0.00	95.62	0.00	41.08	0.00	65.81	0.00	96.90	0.00	22.99	0.00	36.27	0.00	52.74	0.00
Korea Fd	50.56	0.00	81.03	0.00	119.47	0.00	52.44	0.00	84.05	0.00	123.93	0.00	28.12	0.00	45.09	0.00	66.49	0.00
Malaysia Fd	52.17	0.00	83.64	0.00	123.38	0.00	51.74	0.00	82.97	0.00	122.41	0.00	33.61	0.00	53.84	0.00	79.44	0.00
Mexico Eq	35.27	0.00	56.51	0.00	83.25	0.00	37.55	0.00	60.16	0.00	88.63	0.00	20.59	0.00	32.50	0.00	47.30	0.00
Mexico Fd	48.33	0.00	77.45	0.00	114.11	0.00	44.10	0.00	70.65	0.00	104.04	0.00	39.73	0.00	63.58	0.00	93.58	0.00
Taiwan Fd	27.70	0.00	44.34	0.00	65.20	0.00	27.21	0.00	43.55	0.00	64.05	0.00	25.61	0.00	40.72	0.00	59.88	0.00
Turkish Fd	38.87	0.00	62.28	0.00	91.78	0.00	38.89	0.00	62.33	0.00	91.91	0.00	35.81	0.00	57.23	0.00	84.10	0.00

H_0: Series is $I(1)$ (random walk with drift). q is the span of differencing.

M_2 is a standardized test statistic with standard normal distribution and is robust to conditional heteroskedasticity.

Table A7. Estimates of the Fractional Integration Parameter, d (Weekly Country Fund Data).

	Price				*Net Asset Value*				*Cointegrating Residuals*			
	$\hat{d}_{gph}$	*CI*	$\hat{d}_{mle}$	*CI*	$\hat{d}_{gph}$	*CI*	$\hat{d}_{mle}$	*CI*	$\hat{d}_{gph}$	*CI*	$\hat{d}_{mle}$	*CI*
Argentina Fd	0.97	(0.88, 1.05)	0.98	(0.90, 1.05)	1.10	(1.01, 1.19)	1.08	(1.01, 1.15)	0.92	(0.83, 1.01)	0.84	(0.77, 0.91)
Brazil Eq	0.96	(0.88, 1.03)	0.99	(0.94, 1.05)	1.01	(0.93, 1.09)	1.01	(0.95, 1.07)	0.59	(0.51, 0.67)	0.62	(0.57, 0.67)
Brazil Fd	1.02	(0.95, 1.08)	0.97	(0.92, 1.02)	1.01	(0.95, 1.08)	1.04	(0.98, 1.09)	0.72	(0.66, 0.79)	0.73	(0.69, 0.78)
Brazil Inv Trust	0.96	(0.84, 1.09)	1.05	(0.94, 1.16)	0.96	(0.83, 1.08)	0.99	(0.89, 1.09)	0.55	(0.43, 0.68)	0.55	(0.46, 0.64)
Chile Fd	0.95	(0.90, 1.01)	0.97	(0.93, 1.01)	1.08	(1.02, 1.13)	1.06	(1.02, 1.10)	0.80	(0.74, 0.85)	0.80	(0.76, 0.85)
First Philippine Fd	1.00	(0.92, 1.07)	1.00	(0.94, 1.06)	1.10	(1.02, 1.18)	1.11	(1.06, 1.17)	0.72	(0.64, 0.79)	0.74	(0.69, 0.80)
India Fd	1.00	(0.94, 1.06)	1.00	(0.96, 1.05)	1.08	(1.02, 1.14)	1.10	(1.05, 1.15)	0.73	(0.67, 0.79)	0.74	(0.70, 0.78)
India Growth Fd	1.04	(0.97, 1.11)	1.02	(0.96, 1.07)	1.03	(0.96, 1.11)	1.03	(0.97, 1.08)	0.84	(0.77, 0.92)	0.80	(0.75, 0.85)
Indonesia Fd	0.96	(0.90, 1.01)	0.94	(0.90, 0.98)	1.04	(0.99, 1.10)	1.04	(1.00, 1.08)	0.75	(0.69, 0.80)	0.74	(0.70, 0.78)
Jakarta Growth Fd	1.01	(0.93, 1.09)	0.88	(0.82, 0.93)	1.02	(0.94, 1.11)	1.06	(1.00, 1.12)	0.71	(0.63, 0.80)	0.69	(0.63, 0.75)
Korea Eq	1.00	(0.94, 1.06)	0.97	(0.92, 1.01)	1.02	(0.96, 1.08)	0.98	(0.94, 1.03)	0.82	(0.76, 0.88)	0.77	(0.72, 0.81)
Korea Fd	1.05	(1.00, 1.11)	1.02	(0.98, 1.06)	1.08	(1.02, 1.13)	1.05	(1.01, 1.10)	0.85	(0.80, 0.91)	0.77	(0.73, 0.82)
Malaysia Fd	1.00	(0.94, 1.05)	0.96	(0.92, 1.00)	1.07	(1.01, 1.12)	1.06	(1.02, 1.10)	0.78	(0.73, 0.84)	0.77	(0.73, 0.81)
Mexico Eq	0.96	(0.90, 1.02)	0.99	(0.95, 1.04)	1.10	(1.04, 1.16)	1.07	(1.03, 1.12)	0.71	(0.65, 0.77)	0.69	(0.64, 0.73)
Mexico Fd	1.01	(0.95, 1.06)	1.01	(0.97, 1.05)	1.03	(0.98, 1.09)	1.06	(1.02, 1.10)	0.70	(0.65, 0.76)	0.71	(0.67, 0.74)
New South Africa Fd	1.10	(0.98, 1.23)	1.24	(1.12, 1.35)	1.15	(1.03, 1.28)	1.15	(1.04, 1.26)	0.85	(0.72, 0.97)	0.86	(0.76, 0.96)
Taiwan Eq	1.08	(0.96, 1.20)	1.03	(0.93, 1.12)	1.07	(0.96, 1.19)	1.05	(0.95, 1.14)	0.76	(0.65, 0.88)	0.79	(0.71, 0.87)
Taiwan Fd	0.97	(0.92, 1.03)	0.96	(0.92, 1.00)	1.01	(0.95, 1.06)	1.01	(0.97, 1.05)	0.82	(0.76, 0.87)	0.83	(0.79, 0.87)
Thai Capital	0.97	(0.91, 1.02)	0.98	(0.94, 1.02)	1.05	(0.99, 1.11)	1.09	(1.04, 1.13)	0.77	(0.71, 0.83)	0.77	(0.73, 0.81)
Thai Fd	1.00	(0.94, 1.05)	0.98	(0.94, 1.02)	1.04	(0.99, 1.10)	1.06	(1.02, 1.09)	0.81	(0.76, 0.86)	0.82	(0.78, 0.87)
Turkish Fd	1.02	(0.97, 1.08)	1.01	(0.96, 1.05)	1.04	(0.98, 1.09)	1.04	(1.00, 1.08)	0.71	(0.65, 0.77)	0.70	(0.66, 0.74)

$\hat{d}_{gph}$ is the GPH (Geweke and Porter-Hudak, 1983) estimate of the fractional integration parameter.

$\hat{d}_{mle}$ is the Sowell (1992) maximum likelihood estimate of the fractional integration parameter. Estimates are obtained from an ARFIMA$(0, 1 + d, 0)$ model on first differenced data because of the constrained parameter space ($-0.5 < d < .5$).

95% asymptotic confidence intervals in parentheses.

Results from the Reisen (1994) estimator based on the smoothed periodogram and the Robinson (1995) semiparametric estimator yield the same conclusions.

Table A8. Estimates of the Fractional Integration Parameter, d (Daily Country Fund Data).

	Price				*Net Asset Value*				*Cointegrating Residuals*			
	$\hat{d}_{gph}$	*CI*	$\hat{d}_{mle}$	*CI*	$\hat{d}_{gph}$	*CI*	$\hat{d}_{mle}$	*CI*	$\hat{d}_{gph}$	*CI*	$\hat{d}_{mle}$	*CI*
Argentina Fd	1.00	(0.94, 1.07)	1.01	(0.96, 1.06)	1.02	(0.96, 1.08)	1.09	(1.04, 1.14)	0.86	(0.79, 0.92)	0.84	(0.79, 0.88)
Brazil Eq	1.00	(0.95, 1.05)	0.97	(0.94, 1.01)	1.12	(1.07, 1.16)	1.09	(1.05, 1.13)	0.71	(0.66, 0.76)	0.68	(0.64, 0.71)
Brazil Fd	1.00	(0.96, 1.05)	1.03	(1.00, 1.07)	1.02	(0.98, 1.07)	1.04	(1.01, 1.08)	0.72	(0.68, 0.76)	0.71	(0.68, 0.74)
Brazil Inv Trust	0.99	(0.93, 1.04)	1.15	(1.10, 1.20)	1.00	(0.94, 1.05)	1.05	(1.00, 1.09)	0.73	(0.67, 0.78)	0.74	(0.69, 0.79)
Chile Fd	1.03	(1.00, 1.07)	1.03	(1.00, 1.06)	1.06	(1.03, 1.10)	1.06	(1.03, 1.08)	0.87	(0.83, 0.90)	0.85	(0.83, 0.87)
India Fd	1.02	(0.96, 1.07)	1.02	(0.97, 1.06)	1.03	(0.97, 1.09)	1.04	(1.00, 1.09)	0.64	(0.58, 0.69)	0.60	(0.56, 0.64)
Indonesia Fd	1.00	(0.97, 1.03)	0.93	(0.91, 0.95)	1.03	(1.00, 1.06)	1.07	(1.05, 1.10)	0.78	(0.75, 0.81)	0.77	(0.74, 0.79)
Jakarta Growth Fd	0.93	(0.86, 1.00)	0.81	(0.76, 0.86)	1.06	(0.99, 1.12)	1.04	(0.99, 1.10)	0.79	(0.72, 0.86)	0.73	(0.67, 0.78)
Korea Eq	1.00	(0.96, 1.03)	0.94	(0.92, 0.97)	1.01	(0.97, 1.04)	0.99	(0.96, 1.01)	0.51	(0.48, 0.55)	0.50	(0.48, 0.52)
Korea Fd	1.00	(0.97, 1.03)	0.98	(0.96, 1.00)	1.00	(0.97, 1.03)	1.01	(0.99, 1.04)	0.76	(0.73, 0.79)	0.74	(0.72, 0.76)
Malaysia Fd	1.00	(0.97, 1.03)	0.97	(0.95, 0.99)	1.08	(1.06, 1.11)	1.08	(1.06, 1.10)	0.88	(0.85, 0.91)	0.86	(0.84, 0.88)
Mexico Eq	1.03	(0.99, 1.06)	1.02	(0.99, 1.04)	1.04	(1.00, 1.07)	1.01	(0.99, 1.04)	0.73	(0.70, 0.77)	0.72	(0.69, 0.75)
Mexico Fd	1.02	(0.99, 1.05)	1.02	(1.00, 1.05)	1.03	(1.00, 1.06)	1.03	(1.00, 1.05)	0.66	(0.63, 0.69)	0.68	(0.67, 0.70)
Taiwan Fd	1.01	(0.97, 1.05)	0.98	(0.95, 1.00)	1.01	(0.97, 1.05)	0.97	(0.94, 1.00)	0.59	(0.55, 0.63)	0.59	(0.56, 0.62)
Turkish Fd	1.01	(0.97, 1.04)	0.99	(0.96, 1.01)	0.98	(0.95, 1.02)	0.99	(0.96, 1.01)	0.76	(0.72, 0.79)	0.75	(0.73, 0.78)

$\hat{d}_{gph}$ is the GPH (Geweke and Porter-Hudak, 1983) estimate of the fractional integration parameter.

$\hat{d}_{mle}$ is the Sowell (1992) maximum likelihood estimate of the fractional integration parameter. Estimates are obtained from an ARFIMA$(0, 1+d, 0)$ model on first differenced data because of the constrained parameter space $(-0.5 < d < .5)$.

95% asymptotic confidence intervals in parentheses.

Results from the Reisen (1994) estimator based on the smoothed periodogram and the Robinson (1995) semiparametric estimator yield the same conclusions.

Table A9. Fractional error correction model of emerging market closed-end country funds (weekly data, 3-month electoral period).

$$\Delta P_t = \phi_1 + \alpha_1[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{1i}\Delta N_{t-i} + \sum_{i=1}^{L}\omega_{1i}\Delta P_{t-i} + \beta_1 E_t + \upsilon_{1t},$$

$$\Delta N_t = \phi_2 + \alpha_2[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{2i}\Delta P_{t-i} + \sum_{i=1}^{L}\omega_{2i}\Delta N_{t-i} + \beta_2 E_t + \upsilon_{2t}$$

				Long-run adjustment				*Short-run adjustment* H_0: $\gamma' = 0$		*Electoral period*			
Country Fund	N	$\hat{d}$	*Lags*	$\hat{\alpha}_1$	t	$\hat{\alpha}_2$	t	F_1	F_2	$\hat{\beta}_1$	t	$\hat{\beta}_2$	t
Argentina Fd	530	0.84	1	**-0.96**	-3.88	0.08	0.44	1.08	0.05	0.01	1.33	0.01	1.55
Brazil Eq	682	0.62	1	**-0.50**	-4.97	0.10	1.03	**8.05**	0.67	-0.01	-0.72	0.00	-0.10
Brazil Fd	949	0.73	1	**-0.36**	-3.22	**0.23**	2.00	2.82	1.20	0.00	-0.63	0.00	0.39
Brazil Inv Trust	273	0.55	1	**-0.57**	-4.58	0.23	1.85	2.56	2.37	0.00	-0.42	0.00	0.04
Chile Fd	1304	0.80	1	**-0.58**	-3.84	0.04	0.33	1.41	1.48	0.00	-0.34	0.00	-0.55
First Philippine Fd	711	0.74	2	**-1.09**	-5.62	0.21	1.43	**9.57**	**6.65**	0.00	0.56	0.01	0.89
India Fd	1078	0.74	1	**-0.37**	-3.08	0.04	0.40	0.01	0.08	**0.01**	1.88	0.01	1.55
India Growth Fd	772	0.80	1	-0.20	-1.41	**0.46**	3.61	0.01	**4.58**	0.00	-0.18	0.00	-0.01
Indonesia Fd	1280	0.74	3	**-0.93**	-4.78	-0.06	-0.41	**7.10**	**7.13**	**0.01**	1.65	**0.01**	2.18
Jakarta Growth Fd	583	0.69	1	**-0.46**	-3.71	**0.17**	1.99	**5.55**	1.13	**0.05**	2.92	**0.04**	3.14
Korea Eq	1089	0.77	1	-0.23	-1.82	0.18	1.44	0.06	2.93	-0.01	-1.59	**-0.02**	-4.23
Korea Fd	1277	0.77	1	**-0.48**	-2.83	0.05	0.30	1.18	0.50	-0.01	-0.95	**-0.02**	-2.71
Malaysia Fd	1285	0.77	1	**-0.29**	-2.62	**0.25**	3.09	1.69	3.59	0.00	-0.55	0.00	0.65
Mexico Eq	1210	0.69	1	**-0.71**	-6.12	-0.15	-1.46	**8.10**	**4.42**	0.01	0.93	0.00	0.78
Mexico Fd	1380	0.71	1	**-0.32**	-3.24	**0.18**	2.05	1.68	0.11	0.00	0.66	0.01	1.29
New South Africa Fd	275	0.86	1	**-1.18**	-2.40	-0.67	-1.41	1.32	2.16	0.01	1.05	0.00	0.54
Taiwan Eq	303	0.79	1	**-0.66**	-2.31	0.09	0.33	0.44	0.00	0.00	0.25	0.00	-0.52
Taiwan Fd	1397	0.83	1	**-0.35**	-2.94	0.10	1.03	**4.39**	0.03	0.01	1.08	0.00	0.55
Thai Capital	1212	0.77	1	**-0.66**	-5.47	0.04	0.45	**10.91**	0.35	0.00	-0.31	0.00	0.60
Thai Fd	1390	0.82	1	**-0.45**	-3.10	0.15	1.29	**4.26**	0.60	0.00	-0.40	0.00	1.16
Turkish Fd	1295	0.70	1	**-0.48**	-4.85	0.15	1.39	**4.50**	0.01	0.00	0.62	0.00	-0.47

Entries in bold are statistically significant at the 5% level, except for estimates of the election parameter, β, which are reported at the 10% level.

$\hat{d}$ is the Sowell (1992) maximum likelihood estimate of the fractional integration parameter.

Table A10. Fractional error correction model of closed-end country funds with election event study (daily data, 3-month pre-election event window).

$$\Delta P_t = \phi_1 + \alpha_1[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{1i}\Delta N_{t-i} + \sum_{i=1}^{L}\omega_{1i}\Delta P_{t-i} + \sum_{\tau=t_1}^{t_2}\beta_{1,\tau}D_{\tau,t} + \upsilon_{1t},$$

$$\Delta N_t = \phi_2 + \alpha_2[(1-B)^d - (1-B)]z_t + \sum_{i=1}^{L}\gamma_{2i}\Delta P_{t-i} + \sum_{i=1}^{L}\omega_{2i}\Delta N_{t-i} + \sum_{\tau=t_1}^{t_2}\beta_{2,\tau}D_{\tau,t} + \upsilon_{2t}$$

						Long-run adjustment				*Short-run adjustment* H_0: $\gamma' = 0$		*Event study* H_0: $CAR = 0$ ($CAR \equiv \sum \beta_\tau$)			
Country	*Election*	*Fund*	*N*	$\hat{d}$	*Lags*	$\hat{\alpha}_1$	*t*	$\hat{\alpha}_2$	*t*	F_1	F_2	$\widehat{CAR}_1$	χ^2	$\widehat{CAR}_2$	χ^2
Argentina	1999	Argentina Fd	481	0.66	1	**-0.58**	-3.21	0.05	0.35	1.16	0.57	0.04	0.06	0.18	2.48
Brazil	2002	Brazil Eq	520	0.68	2	**-0.61**	-2.97	0.03	0.18	2.90	**3.13**	-0.10	0.39	-0.05	0.11
Brazil	2002	Brazil Fd	520	0.70	2	**-0.58**	-2.25	0.23	0.85	1.03	0.82	-0.14	0.95	-0.05	0.12
Brazil	2014	Brazil Inv Trust	520	0.73	2	**-0.89**	-4.90	-0.04	-0.21	**7.01**	0.45	-0.06	0.72	-0.04	0.36
Chile	2009	Chile Fd	521	0.73	2	0.03	0.13	**0.37**	2.16	**16.74**	**3.30**	0.04	0.06	0.10	0.57
India	2014	India Fd	520	0.60	1	**-0.49**	-3.86	**0.32**	2.30	2.88	1.69	0.12	2.40	0.07	0.65
Indonesia	1999	Jakarta Growth Fd	336	0.72	1	-0.28	-1.43	**0.33**	2.70	0.61	**4.30**	**0.97**	4.69	0.27	0.87
Indonesia	2004	Indonesia Fd	521	0.76	1	**-0.83**	-5.09	-0.02	-0.31	**16.23**	0.30	-0.08	0.07	0.10	0.52
Indonesia	2009	Indonesia Fd	522	0.75	2	**-1.09**	-4.74	0.09	0.49	**12.19**	**3.32**	0.24	1.41	0.23	1.84
Indonesia	2014	Indonesia Fd	522	0.61	1	**-0.69**	-4.46	-0.03	-0.19	3.43	0.25	-0.01	0.00	0.01	0.00
South Korea	1997	Korea Fd	251	0.83	1	0.10	0.28	**1.00**	2.84	0.26	**4.51**	**-0.39**	2.97	**-1.21**	32.96
South Korea	2002	Korea Eq	469	0.36	1	**-0.64**	-8.71	**0.27**	3.68	**21.90**	1.31	-0.02	0.01	-0.22	2.05
South Korea	2002	Korea Fd	522	0.58	1	**-0.74**	-5.93	0.06	0.51	**11.31**	0.14	-0.12	0.91	-0.15	1.47
South Korea	2007	Korea Eq	522	0.66	1	**-0.41**	-4.07	0.11	1.51	1.84	0.73	-0.01	0.01	-0.12	2.07
South Korea	2007	Korea Fd	522	0.71	1	**-0.78**	-3.83	-0.15	-0.78	**5.17**	0.20	**-0.48**	17.71	**-0.50**	21.23
South Korea	2012	Korea Eq	522	0.32	1	**-0.39**	-7.00	**0.36**	4.55	0.11	**11.80**	-0.02	0.08	-0.08	0.74
South Korea	2012	Korea Fd	522	0.47	1	**-0.41**	-3.41	**0.36**	2.81	0.23	**9.76**	0.01	0.01	0.03	0.11
Malaysia	1995	Malaysia Fd	521	0.87	1	**-0.75**	-2.52	-0.10	-0.49	**5.64**	1.05	0.01	0.00	0.17	1.70
Malaysia	1999	Malaysia Fd	520	0.93	1	0.03	0.06	**1.08**	2.65	0.01	**6.26**	-0.24	0.66	-0.02	0.01
Malaysia	2004	Malaysia Fd	521	0.68	1	**-0.60**	-4.58	-0.04	-0.64	3.59	3.76	0.06	0.16	0.09	1.85
Malaysia	2008	Malaysia Fd	522	0.61	1	**-0.55**	-4.28	**0.15**	2.00	1.13	1.16	0.15	1.45	0.00	0.00
Mexico	2000	Mexico Fd	521	0.63	1	**-0.33**	-2.35	**0.46**	3.27	0.41	**5.38**	-0.03	0.04	0.12	0.66
Mexico	2006	Mexico Eq	520	0.80	2	**-1.25**	-3.20	-0.48	-1.39	**6.20**	2.48	-0.03	0.05	-0.02	0.03
Mexico	2006	Mexico Fd	520	0.63	1	-0.26	-1.64	0.13	0.82	0.31	0.03	0.08	1.02	0.01	0.02
Mexico	2012	Mexico Eq	521	0.56	1	**-0.45**	-5.13	**0.34**	3.49	**8.47**	**7.61**	0.07	1.00	0.06	0.64
Mexico	2012	Mexico Fd	521	0.61	2	**-0.52**	-3.22	0.29	1.64	**6.49**	1.71	0.03	0.24	0.10	2.12
Taiwan	2008	Taiwan Fd	522	0.62	1	**-0.33**	-3.13	**0.25**	2.06	1.87	2.51	**0.25**	4.92	0.20	2.35
Taiwan	2012	Taiwan Fd	521	0.48	1	**-0.32**	-5.02	**0.34**	4.64	0.52	**8.54**	**-0.14**	4.62	**-0.16**	4.31
Turkey	2002	Turkish Fd	435	0.67	1	**-0.30**	-2.45	**0.50**	2.72	0.51	1.91	-0.13	0.28	0.12	0.12
Turkey	2007	Turkish Fd	520	0.78	1	**-0.52**	-3.32	0.05	0.29	**6.89**	0.03	-0.05	0.07	-0.04	0.04
Turkey	2011	Turkish Fd	520	0.66	1	**-0.47**	-2.66	0.21	1.16	1.02	1.38	0.02	0.03	0.07	0.46
Turkey	2014	Turkish Fd	520	0.76	1	**-0.76**	-2.01	-0.35	-0.84	0.58	0.06	0.00	0.00	0.01	0.00

Entries in bold are statistically significant at the 5% level, except for *CAR* estimates, which are shown at the 10% level.

$\hat{d}$ is the Sowell (1992) maximum likelihood estimate of the fractional integration parameter.

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