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JÉFFERSON AUGUSTO COLOMBO

ESSAYS IN EMPIRICAL CORPORATE FINANCE AND MACRO-FINANCE

Porto Alegre
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Tese submetida ao Programa de Pós-Graduação em Economia da Faculdade de Ciências Econômicas da UFRGS, como requisito para obtenção do título de Doutor em Economia, com ênfase em Economia Aplicada.

Orientador: Prof. Dr. João Frois Caldeira

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ABSTRACT

In this thesis, I present three empirical essays on corporate finance and macro-finance applied to Brazil. In the first one, I show that an exogenous tax change at the investor level can have real effects on the invested firms' behavior. My evidence suggests that treated firms adjust their financial policies considering substitute financial instruments and seeking to minimize overall tax spending. In the second paper, I analyze the role of equity foreign portfolio investment (EFPI) on affecting aggregate investment. The results show that EFPI has a marginal positive impact on the gross capital formation, but this relation seems to be contingent on institutional factors such as government intervention in credit markets. Finally, in the third essay, I show that an exogenous increase in collateral prices can have positive consequences on firms' financing and investment decisions. The credit expansion registered in Brazil in the middle of the 2000's seem to have alleviated financial constraints most for smaller, less tangible firms, which probably were (at least partially) out of the credit market before the boom.

Keywords: Tax effects. Natural experiment. Interdependence of corporate financial policies. Financial constraints. Equity foreign portfolio investment. Aggregate investment.

RESUMO

Esta tese é composta de três ensaios empíricos sobre finanças corporativas e macrofinanças, todos eles aplicados ao Brasil. O primeiro mostra como uma mudanças tributárias no nível do acionista podem afetar as decisões financeiras das empresas investidas, através da estrutura de propriedade. Os resultados sugerem que as empresas ajustam suas políticas financeiras para minimizar os gastos tributários totais (nível do acionista mais nível da firma). No segundo artigo, analisa-se a relação entre o investimento estrangeiro em carteira (EFPI) e o investimento agregado brasileiro. Os resultados mostram que o EFPI tem um impacto marginal positivo na formação bruta de capital fixo, mas que essa relação é condicionada a fatores institucionais, tal como o grau de intervenção do governo no mercado de crédito. Finalmente, no terceiro ensaio, mostro que um aumento exógeno dos preços dos ativos colateralizáveis imobiliários pode ter consequências positivas no financiamento e investimento das empresas. As firmas aparentemente mais beneficiadas pelo ciclo expansionista de crédito observado no Brasil durante os anos 2000 foram justamente aquelas com menor grau de tangibilidade, potencialmente fora do mercado de crédito no período anterior.

Palavras-chave: Mudanças tributárias. Experimento natural. Interdependência de políticas financeiras. Restricões financeiras. Investimento de carteira de ações no exterior. Investimento agregado.

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1 The Role of Taxes and the Interdependence Among Corporate Financial Policies: Evidence from a Natural Experiment

Abstract

In this paper, we investigate whether and how firms respond to an exogenous tax variation at the investor level by examining their financial decisions following a tax reform for pension funds in Brazil. Consistent with the tax-preference theory of dividends, we find that after implementation of the new law, firms tend to distribute more tax-deductible dividends — called Interest on Equity (IOE) — when the largest or second largest shareholder is a pension fund rather than other types of agents. Surprisingly, control firms also increased (but less than treated firms) their tax-deductible dividend payments, probably to attract more institutional investors and to reduce their cost of capital. We also find that treated firms reduced their leverage relative to control firms after the new law, suggesting that equity tax shields and debt tax shields act as substitute financial instruments. Overall, our evidence suggests that tax is a first-order determinant of corporate financial decisions and firms adjust their policies in consideration of the interdependence among alternative financial instruments.

Keywords: Taxation, Dividend Policy, Interdependence of Corporate Financial Decisions, Natural Experiment.
JEL Codes: G30, G35, G38;

1.1 Introduction

In the Modigliani and Miller (1958) and Miller and Modigliani (1961) frictionless and symmetric information world, corporate financial decisions are irrelevant to firm value. However, when we add market imperfections such as corporate or personal taxes, transaction costs, and asymmetric information, corporate financial decisions become important. Although particular attention has been paid to the influence of taxes on financial decisions, many issues remain unsolved, including whether tax effects are of first-order importance and whether corporate actions are affected by investor-level taxes (GRAHAM, 2003). Fama (2010) also argues that understanding how corporate policies interact with tax incentives is still a significant remaining challenge in corporate finance.

In this paper, we investigate if a taxation shock at the shareholder level causes changes in invested firms' behavior. We also analyze whether these effects are isolated or broader because of the interdependence among firms' financial decisions. These issues are not well understood in the corporate governance and taxation literature, and configure important topics for future research (GRAHAM, 2003; CLAESSENS; YURTOGLU, 2013). By focusing on an exogenous taxation change at the shareholder level rather than at the firm level, our research also relates to large shareholders' ability to influence firms' financial decisions.[1]

One main reason for this gap in the literature is that it is difficult to address credible causal relations among investor-level taxation and firm outputs. First, taxation changes often affect all economic agents, i.e., the treatment is not restricted to a particular group of firms or individuals. Second, the effect passes over the invested firm through the ownership structure, and ownership can be driven by the firms or its shareholders' unobservable characteristics. Third, it is difficult to find institutional environments in which firms can substitute among different tax shields at a relatively low cost. This paper attempts to fill this gap by using a unique setting in Brazil, where corporate tax legislation allows firms to substitute regular, non-deductible dividends with a tax-deductible dividend, called Interest on Equity (IOE).[2] Whereas receiving regular dividends is a tax-exempt

[1] Ownership and control are generally highly concentrated, especially in emerging markets. As pointed out by Claessens et al. (2002), instruments such as cross holdings, pyramidal ownership, and dual classes of shares facilitate tunneling activities (also called the "entrenchment effect").

[2] IOE expenses are tax deductible at the firm level in the same manner as debt. Henceforth, considering a statutory corporate tax of 34% in Brazil, each $1 paid out in the form of IOE instead of regular dividends can reduce firms' earnings before taxes by up to $0.34.

event for all shareholders, receiving IOE is taxed at 15%, which eliminates part of the tax benefits at the firm level.

The empirical strategy of this paper consists of exploiting an exogenous variation in the tax rate of IOE receipts to estimate its impacts on corporate behavior. Given the enactment of the law 11,053/2004, the "pension funds tax reform", all pension funds receive tax exemption on IOE income, whereas all other shareholders continued to be taxed at 15%. This legal change created a quasi-natural experiment in the Brazilian stock market that allows us to estimate the causal effect of an investor-level taxation change on a firm's dividend policy and debt policy.

Specifically, we use the differences-in-differences (DD) estimation to identify whether firms with pension funds as the first or second largest shareholder tend to use more IOE over its total payout relative to similar firms with lower or no pension fund ownership. We also use a propensity score matching approach to run this regression for both matched and non-matched samples. This identification strategy has become increasingly common in the empirical corporate finance literature, particularly to estimate a causal relationship between the financial environment and corporate decisions (see, for example, ALMEIDA et al., 2012; LEMMON; ROBERTS, 2010; MICHAELY; ROBERTS, 2012; GARCIA-APPENDINI, 2015).

The DD approach allows us to control for both unobservable heterogeneity and a potential selection bias regarding pension fund ownership. We also control for variables that can affect payout decisions, such as size, profitability, leverage, liquidity, ownership concentration, and separation of cash flow and control rights. Finally, we use industry fixed effects to control for specific industry shocks and year fixed effects to control for common economic shocks.

We argue that Brazil configures a nearly ideal setting to answer our research question. First, ownership concentration is particularly high in Latin America, and especially in Brazil. According to Claessens and Yurtoglu (2013), the typical largest shareholder owns more than 50% of the voting shares in Latin America, and more than 60% in Argentina and Brazil. In addition, non-voting stock and dual-class shares are more prevalent in Latin American than in other emerging regions such as East Asia. Second, we observe in the Brazilian capital market a large variation in the legal nature of controlling shareholders and explore this cross-sectional variation to identify the invested firms that are more likely to be affected by the pension fund reform that we study. Third, Latin America has

the most time-consuming tax system of the world, and the average results for Brazil are even worse than the average for America and the Caribbean (WORLD BANK, 2015). Therefore, we expect that any regulatory change that could reduce the overall tax burden for shareholders should be strongly pursued. Finally, the presence of an equity tax shield instrument (IOE) that is a substitute for debt makes the relation between dividend policy and capital structure even more complex and sensible to investor-level tax changes.

Our first results regard the composition of firms' payouts. Consistent with the tax preference of dividends, we find that 76.8% of treated firms paid out IOE after the reform compared with an average of 17.7% during the pre-treatment period. The DD estimate confirms that the differences are not only statistically but also economically significant. The fraction of firms that use IOE increased 21 percentage points (p.p.) more in the treated than in the control group. We also estimate a quantitative effect on continuum variables: we estimate a 12.9 p.p. larger increase on IOE / Total Payout for the treated group, which represents a sharp increase of 112.2% considering the pre-event mean of 11.5%.

Somehow surprisingly, the control firms of our sample, which are similar in several observable dimensions but do not have pension funds as first or second largest shareholders on voting shares, also increased (but far less than treated firms) their IOE payments after the new law. Although the policy change does not create immediate tax gains to their controlling shareholders, managers of these firms could set higher IOE payments to attract more institutional investors (clientele effect, as in BECKER; IVKOVIC; WEISBENNER, 2011) and, therefore, to reduce the weighted-average cost of capital.

Our second analysis focuses on the consequences of the new law to other firms' financial decisions. We find a negative and statistically significant relation between IOE payments and Debt / Total Assets, especially after the tax reform. Although treated firms increased their IOE cash payments for tax reasons after the pension fund tax reform, these firms also reduced their leverage relative to control firms by 2.6 p.p. (or 5.2%, considering the pre-treatment leverage average). This empirical evidence suggests that equity tax shields (ETS) and debt tax shields (DTS) are substitutes, i.e., firms increase (decrease) the use of less (more) costly tax shields. Finally, our evidence suggests that firms jointly determine dividend and capital structure policies. This evidence of in-

terdependence among corporate financial decisions is consistent with other theoretical and empirical findings (JENSEN; SOLBERG; ZORN, 1992; GIVOLY et al., 1992; GATCHEV; PULVINO; TARHAN, 2010; LIN; FLANNERY, 2013) and suggests that tax reforms can have broader effects on the economy through indirect channels.

Overall, the results of this paper can be related to different topics in the literature. First, this paper provides evidence that taxes can be a first-order determinant for firms' financial decisions, which answers one of the unresolved issues related to taxes as exposed in Graham (2003). This evidence also contrasts that found in Brav et al. (2005), who argues that taxes play a secondary role in corporate finance. The results of their survey of CFOs in the United States suggest that historical dividend levels are first-order determinants of dividend choices. However, our evidence that firms react sharply to investor-level changes in taxation is consistent with recent empirical evidence from countries such as the United States (BRAV; JIANG; KIM, 2015) and Canada (DOIDGE; DYCK, 2015).

Second, our results also suggest that regulatory changes at the shareholder level can be transposed to invested firms, especially in markets with high ownership concentration. Considering our evidence of interdependence among financial decisions, the effect of tax reforms can be wider than expected because firms experience not only direct but also important indirect effects. From the point of view of policy makers, our results suggest that the enactment of new laws and tax reforms can generate multiplier effects given the interdependence between ownership structure and firms' outcomes.

Finally, our results provide a better understanding of the reasons why such a large number of companies in Brazil prefer to pay out their earnings as regular dividends rather than IOE. Although some progress has been made (see, for example, BOULTON; BRAGA-ALVES; SHASTRI, 2012), the reasons why such a large number of firms choose to distribute only non-deductible dividends is still unclear. Our empirical evidence suggests that ownership structure and taxation at the investor level can explain a considerable part of this puzzle. If the sum of taxation at both the firm and the investor levels is higher than the tax benefits of IOE distributions, than it is perfectly natural for firms to prefer to use regular, non-deductible dividends. The sharp increase in IOE payments after the pension fund reform — which reduced the cost of receiving IOE payments — corroborates this

hypothesis.

The remainder of the paper is structured as follows. Section 2 describes the institutional environment in Brazil, with a focus on IOE distributions and tax considerations. We also list and compare instruments similar to IOE that are used around the world. In section 3, we outline our empirical strategy and data collection process. In section 4 and 5, we present the results of the paper and a discussion about its main implications, respectively. Finally, in section 6, we present the conclusions of this paper.

1.2 Characteristics of the Brazilian Capital Market

The Brazilian capital market has some essential singularities. In this section, we highlight the high ownership concentration, the existence of IOE payout distributions and the corporate costs derived from both size and complexity of the country's tax system.

1.2.1 High ownership concentration

For Brazil, similar to most countries whose legal system descends from French civil law, concentration is a fundamental characteristic of the country's ownership structure (LA PORTA et al., 2000). In comparative terms, the Brazilian market is closer to those of Japan and Continental Europe, and less close to the markets of the United States and the United Kingdom (CANELLAS; LEAL, 2009). Nevertheless, ownership concentration has been changing over time. During the last decade, Brazil has faced a wave of corporate restructurings caused by privatization and the entry of new partners of private sector companies, notably foreign and institutional investors (SILVA, 2004).

A major change in corporate law in the Brazilian corporate environment occurred with the entry into force of Law 10.303/2001, also known as the New Corporate Law. Previously, legislation allowed companies to issue up to two-thirds of total capital in the form of shares without voting rights (preferred shares). Ultimately, a company might exercise majority control with only 16.67% of total capital, which gave rise to misaligned management practices in terms of risk and returns on capital. With the introduction of the New Corporate Law, the proportion of common and preferred shares fell from two-thirds to 50%, but only for new public companies. Canellas and Leal

(2009) suggest that firms that went public after 2001 present greater dispersions in their control structures. Black, Carvalho and Gorga (2010) show that a high percentage of Brazilian privately controlled firms (84%) issue non-voting preferred shares, revealing that this practice is common in the Brazilian capital market. In terms of value, Black, Carvalho and Gorga (2012) find empirical evidence that the voting-to-common shares ratio is directly related to Tobin's Q (a proxy for firm value) in Brazil. This result seems consistent with those presented by Claessens et al. (2002): in East Asia, firm value generally increases with the share of cash flow rights owned by the largest shareholder and decreases with the separation between control and cash flow rights.

1.2.2 Interest on Equity (IOE)

Given the end of automatic monetary correction, Law 9249/1995, which introduced the concept of IOE, came into effect as of January 1, 1996. Article 9, Paragraph 7 of this legislation allows companies to impute interest paid as remuneration of equity to the value of mandatory dividends specified in the Corporations Law. Beginning the following year, 1997, the total amount of interest paid as remuneration of equity had to be limited to a maximum of half the computed earnings before deduction of interest or accumulated profits and profit reserves. This change is in accordance with the provisions of Article 79 of Law 9430/1996.

In short, the IOE institution represents a tax incentive for capital, parallel to the previously existing tax benefit for debt. Incidentally, debt is widely used in other parts of the world. Allowances for corporate equity (ACE) are found in countries such as Brazil, Italy, and Belgium, making the internal environment for dividend policies even more complex and peculiar in these economies.

Regarding the legal interpretation and despite receiving the name "interest", IOE is more similar to dividends than to interest. Resolution 207/96 of the Comissão de Valores Mobiliários (CVM) states that, regarding the concept of profit in corporate law, the distribution of returns on equity constitutes a distribution of income and not expenditure. Moreover, the regulatory organ affirms that if such interest is not treated as a distribution of income, the comparability of public company results will be affected. Therefore, repercussions may occur in all holdings and allocations are calculated on the basis of corporate profit.

1.2.3 A general view of corporate taxation in Brazil

Brazilian companies face significant taxes and contributions. However, more impressive is the shocking complexity of taxes and contributions. According to the World Bank (2015), an average firm located in Sao Paolo — the financial capital of the country — spends 2,600 hours per year paying its taxes (Table 1). This figure is high relative to the OECD average (175.4 hours) and the Latin America and Caribbean average (365.8 hours). In all dimensions covered by the Doing Business annual report (WORLD BANK, 2015), the worse position Brazil occupies is precisely paying taxes (177th out of 188 countries).

Table 1: Paying Taxes in Brazil

Indicator	Brazil (São Paulo)	Latin America and Caribbean	OECD
Payments (number per year)	9.0	29.9	11.8
Time (hours per year)	2.600	365.8	175.4
Profit tax (%)	24.8	20.7	16.4
Labor tax and contributions (%)	40.3	14.7	23.0
Other taxes (%)	3.8	12.9	1.9
Total tax rate (% profit)	68.9	48.3	41.3

Source: Data from Doing Business 2015 (WORLD BANK, 2015).

In addition to time costs, the complexity of corporate taxes also imposes large financial costs to firms. The average company that trades shares on the Brazilian Stock Market must pay income tax (15%), additional income tax (10%), and make a social contribution on net income (9%), as well as PIS and COFINS (9.25%) on total revenue. Together, the tax burden for a company may exceed 43.25% of profits before income taxes. In this context of costly corporate taxes, we expect that firms react optimally to the introduction of any legal tax planning instrument available, such as the one created by the pension tax reform studied in this paper.

1.3 Data and Empirical Strategy

In this subsection, we describe the characteristics of our data, the "natural experiment", and the details of our empirical strategy.

1.3.1 Data

We use financial and ownership structure data from Economatica software for the period 2002–2008. Our initial sample covers all publicly traded Brazilian firms. Our period covers three years before to four years after the regulatory change that became effective on January 1, 2005. We also collect additional information about the companies in our sample through the CVM, the Annual Information Report (IAN), and the Reference Form (FR). This additional data give us detailed information on the companies' main shareholders, in particular to observe the indirect ownership structures.

Using this starting dataset, we dropped additional classes of shares in a given company, i.e., we kept in the sample only one class of share per firm.[3] Because our variables are essentially at the firm level, keeping more than one observation per firm-year simply duplicates our data for some firms and, therefore, creates a bias in our sample. We also exclude finance industry firms given their peculiarities related to capital structure. Finally, because the analysis aims to understand the distribution of earnings (the choice between dividends and IOE), we deleted observations with a dividend per share of zero and firms that did not have available financial information for at least six years during the 2002–2008 period.[4] The final sample included 108 firms and 636 firm-year observations. To control for the influence of outliers, financial, ownership, and control structure continuous variables were winsorized in each tail at the 5% level. All variables used in this study are described in Table 2.

[3] We kept in our sample only the class of shares with higher trading volume during the entire period. However, because our data cover firm-level characteristics, the method used to drop excess shares in a given firm would not produce different results.

[4] Because we want to analyze the changes before and after the new legislation, we had to keep in the final sample only firms that had sufficient information in both the pre- and post-event periods.

Table 2: Description of variables

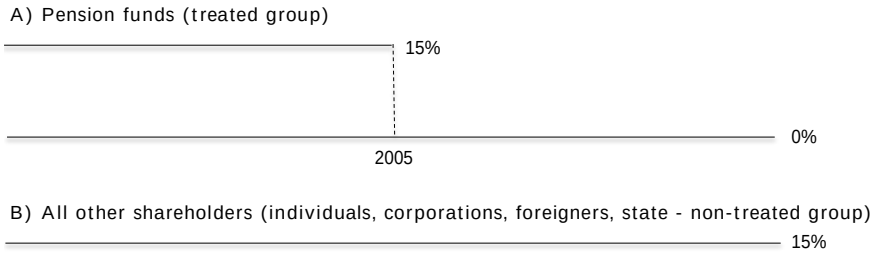
Variable	Description
IOE Dummy	Dummy equal to “1” if the firm distributed IOE in the current year, and "0" otherwise
IOE / Total Payout	Ratio between the IOE amount and total earnings distributed in cash in the current year
IOE / IOE*	Ratio between the IOE amount and the maximum allowed by law in the current year
Ln (Total Assets)	Natural logarithm of Total Assets
Debt / Total Assets	Ratio between the book value of Debt and Total Assets
EBIT / Total Assets	Ratio between EBIT and Total Assets in the current year
Good Governance	Dummy equal to “1” if the firm is listed at any level of the Corporate Governance special segments of BMandF Bovespa
ADR	Dummy equal to “1” if the firm has ADRs on the NYSE, and "0" otherwise.
Current Liquidity Ratio	Ratio between Current Assets and Current Liabilities
Investment / Total Assets	Ratio between Investment and Total Assets
% Cash Flow Largest Shareholder	% of a firm’s cash flow held by the largest shareholder
(% Cash Flow Largest Shareholder)^2	Square of the % of a firm’s cash flow held by the largest shareholder
% Excess Cash Flow Largest Shareholder	% of excess cash flow relative to the voting power held by the largest shareholder
(% Excess Cash Flow Largest Shareholder)^2	Square of the % of excess cash flow relative to the voting power held by the largest shareholder

Source: Authors’ elaboration using data from the Economatica database.

1.3.2 The "Natural Experiment" and the evolution of IOE payments

At the end of 2004, the Federal Government of Brazil signed law number 11,053/2004, legislation that changes the taxation of both private and public pension funds and insurance companies. As of January 1, 2005, pension funds were exempt from taxes on their investment earnings and income as long as the funding came from the participants or the assisted. This law represents a significant change in the tax environment of security companies and pension funds; prior to this law, such entities were taxed at the fund level and beneficiaries were taxed again at the investor level (double taxation). Additionally, the new legislation made it possible for participants to deduct the value of their contributions from the calculation basis for income tax, creating an additional incentive for the capitalization of those firms. Figure 1 shows the basic identification strategy of our paper.

Figure 1: Identification Strategy: Interest on Equity Taxation for Different Shareholders



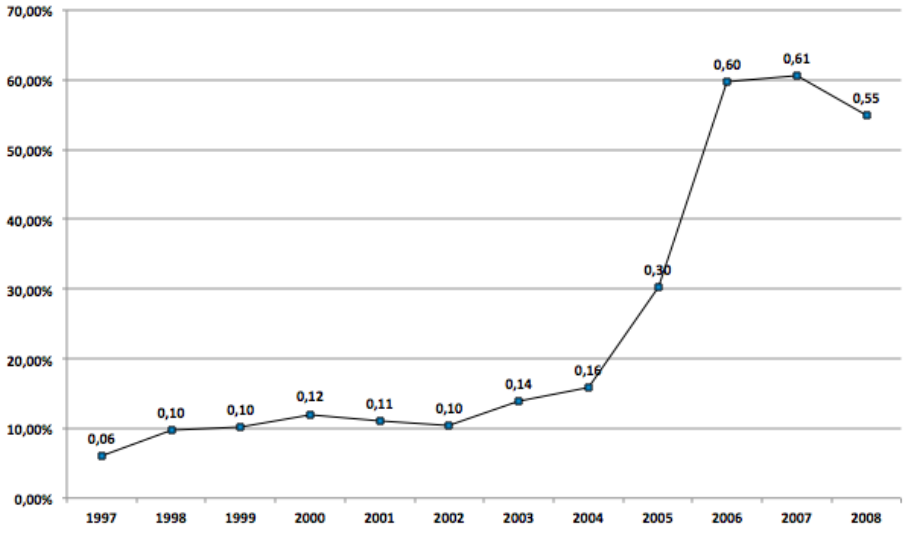
Source: Authors' elaboration.

Our first evidence from our data refers to the evolution of IOE payments in our sample. Using data on IOE and dividend distributions, we (by year) calculate the percentage of eligible firms that distributed any amount of IOE to shareholders in a given fiscal year.[5] The results are exposed in Figure 2. As this figure shows, a sharp increase in the use of IOE for dividend payer firms appears in 2005. The average number of listed firms (private or public) that distributed cash using IOE increased from approximately 16% to 60% in two years. This structural change coincides with the entry into force of the so-called pension fund reform — law number 11,053/2004. Our hypothesis is that the tax changes for associative investment entities as dictated by the new law may

[5] Not all firms are eligible to pay out earnings in the form of an IOE. From this calculation, we delete firms that simultaneously satisfy the following two conditions: i) earnings before interest on equity less than or equal to zero, ii) accumulated earnings plus reserves equal to or less than zero.

have contributed significantly to the sudden increase in the number of companies that have paid IOE precisely through the high concentration and the significant power that large shareholders have in Brazil. Figure 2 shows the basic motivation of this paper and presents an open research question: What are the factors that drive firms to utilize less or more tax-deductible dividends (IOE) instead of regular, non-deductible dividends?

Figure 2: Percentage of eligible firms that distributed cash payouts in Brazil by IOE, 1997–2008



Source: Authors' elaboration.

1.3.3 Empirical Design

Estimating the influence of the controlling shareholder over a firm's dividend decision is difficult for certain causes. Even among companies that are similar in several dimensions, we may observe that ownership concentration is higher for unobservable reasons. In addition, portfolio firms managed by activist investors (such as pension funds, hedge funds, venture capitalists, and others) are likely to differ along other dimensions, such as better future prospects. This phenomenon is known in the literature as selection bias and is the result of better ability to select invested firms (GIROUD; MUELLER, 2015; BRAV; JIANG; KIM, 2015). We overcome this endogeneity issue using a strategy similar to that in Bernstein, Giroud and Towsend (2016) and Giroud and Mueller (2015): instead of attempting to randomize the sample of firms, we approximate our study to an ideal experimental setting by randomly changing the dividend policy after companies

are selected. By doing this, we can estimate the effect of the controlling shareholder on firms' decisions, holding company selection fixed.

We also address some concerns regarding the differences between treated and non-treated firms. Although both groups seem similar in a large number of observable dimensions (see Table 2), we perform a Propensity Score Matching to check the robustness of our results. We discuss next this procedure and the DD approach, which is the basic empirical strategy of this paper.

1.3.3.1 Propensity Score Matching

We construct a matched sample on the basis of a firm's characteristics, such as size, profitability, non-ETS, corporate governance, ownership concentration, and excess control. Intuitively, we must ensure that the association between the tax reform and a firm's IOE payments is actually led by the reform (and not for other firm characteristics that can be correlated with the treatment). More specifically, if firms that have pension funds with a relevant stake in its ownership structure have characteristics that differ from the others, then the estimated effect can be noisy. To address this problem, the corporate finance literature has broadly used the propensity score match (see, for example, LEMMON; ROBERTS, 2010 and MICHAELY; ROBERTS, 2012). Intuitively, the objective is to control for firm characteristics related to pension fund ownership, such as size, leverage, industry, and so on. Therefore, we estimate the following equation:

$$\mathsf{Treat}_i = \beta_0 + \beta_1 \cdot \mathsf{size}_{it} + \beta_2 \cdot \mathsf{leverage}_{it} + \beta_3 \cdot \mathsf{industry}_{it} + \epsilon_{it}, \tag{1}$$

where Treat_{it} is equal to one if firm i has a pension fund as one of the two largest shareholders in the year before the reform, and zero otherwise. Propensity score matching creates a pseudo-random sub-sample in which firms with similar characteristics differ by receiving (treatment group) and not receiving (control group) the treatment. The model is estimated using a *logit* regression.

We estimate the previous equation separately for each year such that each observation has its own fitted value. We use this yearly fitted value to match each treated observation to a non-pension fund firm with the closest fitted value. We also require that the matched firms share the same industry code.[6] Because the non-treated sample is similar to the treated sample in a significant

[6] This requirement is imposed to control for economic shocks that affect more or less one particular industry. If a shock occurs in any particular year, we expect that it will similarly affect all matched firms, i.e., it will not affect our

number of observable dimensions, we use our matched sample only as a robustness check. The results remain the same.

1.3.3.2 Difference in Differences (DD) Approach

The basic empirical strategy of this paper is to use a DD setting. The regression we estimate is as follows:

$$Y_{it} = \delta_0 + \delta_1 \cdot \mathsf{post}_t + \delta_2 \cdot \mathsf{treated}_i + \delta_3 \cdot (\mathsf{treated}_i \times \mathsf{post}_t) + \epsilon_{it}, \tag{2}$$

where i, g, and t represent index firms, industry, and year, respectively; Y_{igt} is the outcome of interest (we use three different dependent variables related to IOE and dividend payments); δ_1 is the estimate of the aggregate factors that cause changes in Y even in the absence of a policy change; δ_2 is the estimate of the differences between the treatment and control groups prior to the policy change; δ_3 is the coefficient of interest; and ϵ_{it} is the white noise error term. Estimating the coefficient is equivalent to calculating the following equation:

$$\widehat{\delta_3} = \left(\widehat{y}_{(\mathsf{treat},\ \mathsf{post}=1)} - \widehat{y}_{(\mathsf{treat},\ \mathsf{post}=0)}\right) - \left(\widehat{y}_{(\mathsf{control},\ \mathsf{post}=1)} - \widehat{y}_{(\mathsf{control},\ \mathsf{post}=0)}\right). \tag{3}$$

According to Imbens and Wooldridge (2009), inference based on even moderate sample sizes in each of the four groups is straightforward, and it is also easily made robust to different group/period variances in the regression framework.

In addition to the basic strategy described in equation (2), we also follow Bertrand, Duflo and Mullainathan (2004) and incorporate a vector of covariates (X) into equation (10), as well as robust and clustered standard errors at the firm level.[7] The general model considered is:

$$Y_{igt} = \alpha + \delta_3 \cdot (\mathsf{treated}_i \times \mathsf{post}_t) + \beta \cdot X_{igt} + \gamma_g + \lambda_t + \epsilon_{it}, \tag{4}$$

where i, g, and t represent index firms, industry, and year, respectively; X_{igt} is a vector of covariates; γ_g are industry fixed effects; λ_t are year fixed effects; and ϵ_{it} is the error term. Analogous to

results.

[7] Reasons for including covariates include efficiency, checks for randomization, and adjustments for conditional randomization (ROBERTS; WHITED, 2013).

the standard DD equation described in (10), the treatment effect is given by δ_3.

1.4 Results

We now analyze the main results of our tests. We start showing the descriptive statistics, and then we analyze the characteristics of our treatment and control group. Following this initial analysis, we show the results of our different empirical approaches.

1.4.1 Descriptive Analysis

We start by presenting the descriptive statistics of the variables used in this study (Table 3). During the analyzed period (2002–2008), the average of the IOE dummy variable is 0.40, indicating that approximately 60% of the firm-year cash distributions in the sample were made exclusively through dividends and only 40% paid out some amount as IOE. Furthermore, the average variable IOE / Total Payout (0.259) shows that, from all of the cash payouts in our sample, only 26% of the total amount was distributed as IOE (complementarily, 74% of the average cash distribution was paid out as regular dividends). The average of the variable IOE/IOE* (0.253) implies that the average firm in our sample distributed only $\frac{1}{4}$ of the total IOE payment allowed by fiscal legislation. Despite the tax advantages at the firm level, we conclude that IOE was not used by a large number of companies during the period.

Table 3 also shows that the average firm in our sample has R$7.8 billion (US$3.9 billions) in total assets, 53% of Total Debt / Total Assets, 11% of EBIT / Total Assets, 22% are committed to one of the Novo Mercado differentiate levels of Corporate Governance, 14% negotiate ADR shares on the NYSE or NASDAQ, 1.79 is the average Current Liquidity Ratio, and 68% is the average ratio of Investments to Total Assets. We also document the descriptive statistics for the ownership concentration variables: the average largest shareholder owns 43% of the firm's total cash flow, and the square of this measure equals 35%.[8] The average excess control of the largest shareholder is 18 percentage points (p.p.), i.e., on average the largest shareholder holds 18 p.p. more voting power than the firm's cash flow.[9] The square of this variable averages 6 p.p.

[8] We include the square of the ownership concentration variables in our analysis to control for a potential non-linear association between ownership concentration and firm outputs. A similar procedure is used by Claessens et al., 2002.

[9] This concept refers to the classic separation between control rights and cash flow rights, as suggested by La Porta et al. (2000). Some of the common mechanisms to leverage control over firms' decisions are pyramid ownership,

Table 3: Descriptive statistics of the final sample, 2002–2008

Note: This table reports the descriptive statistics of our final sample for the 2002–2008 period. The data come from the Economatica Database and initially comprise all Brazilian publicly traded firms. We use the following abbreviations: SD = Standard Deviation, Min = Minimum, Max = Maximum, and N = Number of observations.

Variables	Mean	Median	SD	Min	Max	N
IOE Dummy	0.402	0.000	0.491	0.000	1.000	682
IOE / Total Payout	0.259	0.000	0.377	0.000	1.000	682
IOE / IOE*	0.253	0.000	0.378	0.000	1.000	724
Ln (Total Assets)	14.65	14.85	1.792	10.80	17.29	718
Debt / Total Assets	0.526	0.532	0.164	0.207	1.046	718
EBIT / Total Assets	0.111	0.104	0.070	−0.095	0.236	718
Good Governance	0.218	0.000	0.413	0.000	1.000	724
ADR	0.144	0.000	0.351	0.000	1.000	724
Current Ratio	1.788	1.625	0.843	0.160	4.030	718
Investment / Total Assets	0.677	0.713	0.174	0.072	0.936	718
% Cash Flow Largest Shareholder	0.427	0.387	0.232	0.045	0.981	636
(% Cash Flow Largest Shareholder)^2	0.346	0.272	0.263	0.006	0.965	636
% Excess Cash Flow Largest Shareholder	0.178	0.149	0.173	−0.303	0.664	636
(% Excess Cash Flow Largest Shareholder)^2	0.062	0.024	0.088	0.000	0.441	636

Source: Authors' elaboration.

1.4.2 Are there ex-ante observable differences between treated and non-treated firms?

We follow our analysis by splitting our sample in two periods, i.e., pre and post-pension fund reform. The first period goes from 2002 to 2004 and the second from 2005 to 2008. We also separate firms on the basis of treated (those with relevant participation of pension funds in the sample) and non-treated (others in the final sample). Our first question is: Are there (observable) differences in firms' characteristics between these two groups *before* treatment? Table 4 provides evidence to answer this question.

Even without the matching procedure, the resulting full sample (Panel A) appears to have a good balance between non-treated and treated firms for a large number of dimensions. We focus our attention on the independent variables; however, the dependent variables (the first three in Table

cross-listings, and dual share classes.

Table 4: Ex-ante difference of means for treated and non-treated firms, 2002–2004

Note: This table reports the descriptive statistics of our final sample for the 2002–2008 period. The data come from the Economatica Database and initially comprises all Brazilian publicly traded firms. We use the following abbreviations: SD = Standard Deviation, Min = Minimum, Max = Maximum, and N = Number of observations.

Variables	Non Treated (1)		Treated (2)		Diff. (1 - 2)
	n	Mean	n	Mean	Mean Diff.
Panel A: Full Sample					
IOE Dummy	229	0.153	62	0.177	-0.025
IOE / Total Payout	229	0.103	62	0.115	-0.012
IOE / IOE*	244	0.099	65	0.071	0.027
ln (Total Assets)	244	14.54	65	14.71	-0.169
Debt / Total Assets	244	0.526	65	0.502	0.024
EBIT / Total Assets	244	0.116	65	0.121	-0.006
Good Governance	244	0.139	65	0.169	-0.030
ADR	244	0.131	65	0.138	-0.007
Current Ratio	244	1.684	65	1.877	-0.194
Investment / Total Assets	244	0.691	65	0.700	-0.009
(% Cash Flow Largest Shareholder)^2	209	0.333	54	0.301	0.032
% Excess Cash Flow Largest Shareholder	209	0.168	54	0.226	-0.057^{**}
(% Excess Cash Flow Largest Shareholder)^2	209	0.054	54	0.086	-0.032^{**}
Panel B: Matched Sample (after the Propensity Score Matching)					
IOE Dummy	105	0.124	52	0.212	-0.088
IOE / Total Payout	105	0.087	52	0.137	-0.050
IOE / IOE*	105	0.074	52	0.089	$-0,015$
Ln (Total Assets)	105	14.18	52	14.41	$-0,233$
Debt / Total Assets	105	0.496	52	0.496	$-0,001$
EBIT / Total Assets	105	0.108	52	0.118	$-0,009$
Good Governance	105	0.133	52	0.212	-0.078
ADR	105	0.143	52	0.173	-0.030
Current Ratio	105	1.809	52	1.913	-0.104
Investment / Total Assets	105	0.690	52	0.704	-0.015
% Cash Flow Largest Shareholder	105	0.405	52	0.404	0.001
(% Cash Flow Largest Shareholder)^2	105	0.306	52	0.311	-0.005
% Excess Cash Flow Largest Shareholder	105	0.188	52	0.229	-0.040
(% Excess Cash Flow Largest Shareholder)^2	105	0.063	52	0.088	-0.025

Source: Authors' elaboration.

4) also seem balanced during the pre-event period.[10] We find that size, leverage, profitability, corporate governance, liquidity, investment, and participation on cash flow by the largest shareholder are not statistically different between the two groups. The only exception from Panel A is excess cash flow held by the largest shareholder, which is 5.7 p.p. higher on average for the treated group.

Alternatively, we show in Panel B the comparison between control (matched) and treated groups. We conduct Propensity Score Matching as described in 3.3.1. As the table shows, the statistically significant differences between groups disappear after the matching procedure. However, this disappearance comes with a cost: the number of observations decreases from more than 200 to 105 before the event. We also eliminate some of the treated observations, more precisely those with a missing independent variable or out of the common support region.[11] Following Lemmon and Roberts (2010), we use a large proportion of non-treated firms and admit up to four controls for treated observations, thus admitting replacement.[12]

A more detailed analysis of the p-score differences reveals that the results are very similar to those of other studies using the same technique. For example, Lemmon and Roberts (2010) report that the maximum difference in p-scores between treated and control observations is 0.04, similar to our results.[13] Moreover, our means and medians for the variable are very similar to those found by the referred authors — both are 0.00. Overall, these results indicate that the matching process is accurate and the control firms are similar to the treated firms in several dimensions.

1.4.3 Do firms controlled by pension funds increased their IOE payments after the reform?

In this subsection, we ask whether the treated firms responded differently to the exogenous shock that affected the taxation on IOE receipts for pension funds. To answer that, we divide our analysis in graphical evidence, difference-in-differences analysis, evidence from multivariate models and falsification tests.

[10] According to the literature, pre-event differences in levels on the outcome variable are not a significant problem. However, these variables' trends must be similar to assure the assumption often called "parallel trend" in DID analyses.

[11] The common support option imposes the condition that only the propensity scores in between the lower score from the treatment group and the highest score from the control group is used. In other words, firms with too different propensity scores are not paired together.

[12] One control observation can be matched for more than one treated observation if the propensity scores are close enough.

[13] Our maximum difference in the p-score variable is 0.043.

1.4.3.1 Graphical evidence

We first begin with a graphical analysis. First, in DD analysis, it is important to ensure that both treated and control (or non-treated) groups had the same path for the outcome variables before the event — the "parallel trend" assumption.[14] Figure 16 shows the mean (left side) and the median (right side) of all three IOE variables (IOE Dummy, IOE / Total Payout, and IOE / IOE*) for the treated and non-treated groups by year. We observe from the figure that the parallel trend is very likely to hold for all three different outcome variables. The graphic also shows the visual effect of the pension fund reforms for different firms — beginning in 2005, firms that had significant pension fund participation regarding voting rights (treated firms) increased their IOE payments by a lot more than other firms (non-treated). This increase in the use of IOE immediately after taxation at the investor level declined from 15% to 0% for pension funds, indicating two important points: changes in taxation at the investor level can lead to changes in invested firms' cash payout decisions, and controlling shareholders can exert considerable influence over firms' financial decisions.

A closer look at Figure 16 reveals that the post-event results do not seem to be driven by trends in the outcome variables during the pre-event period (2002–2004). Before the Law, 17.7% of treated firms and 15.3% of non-treated firms made at least one IOE distribution during a given year (IOE Dummy). In addition to the propensity of payments, the level of each annual IOE distribution also seems similar: an average of 11.5% (10.3%) of the total payout was distributed through IOE for the treatment (non-treatment) group, and 7.1% (9.8%) of the maximum annual IOE payments allowed by the legislation were made by the treatment (non-treatment) groups. Based on these years path, it is likely that the sharp increase in IOE payments observed in the treated group would not have occurred in the absence of the pension fund reform.

1.4.3.2 Difference-in-Differences (DD) analysis

Following the graphical analysis, we now provide evidence for our basic empirical strategy: the DD analysis. We estimate equation (2) using our three different measures of IOE payments as dependent variables. The results are reported in Table 5.

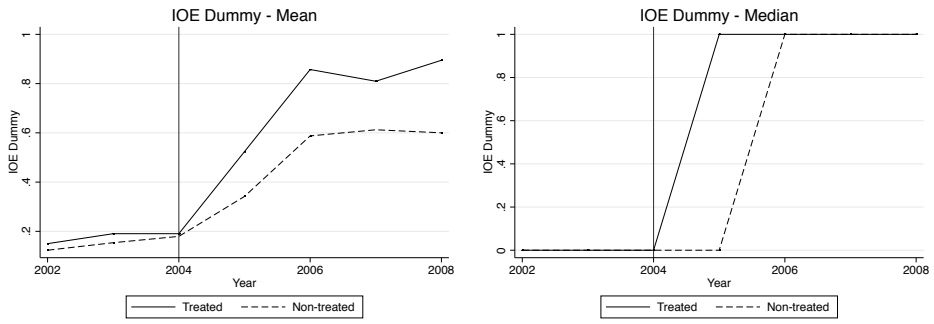
Basically, the results reported in Table 5 confirm the previous suggestions made using the

[14] The parallel trend assumption requires that both groups must have similar paths before the event and not necessarily at the same levels (LEMMON; ROBERTS, 2010). If the parallel trend is not satisfied, eventual post-event differences in outcomes could have been observed even in the absence of the treatment (i.e., it could lead to spurious inferences).

Figure 3: Mean and median of all IOE variables for treated and non-treated firms, 2002–2008

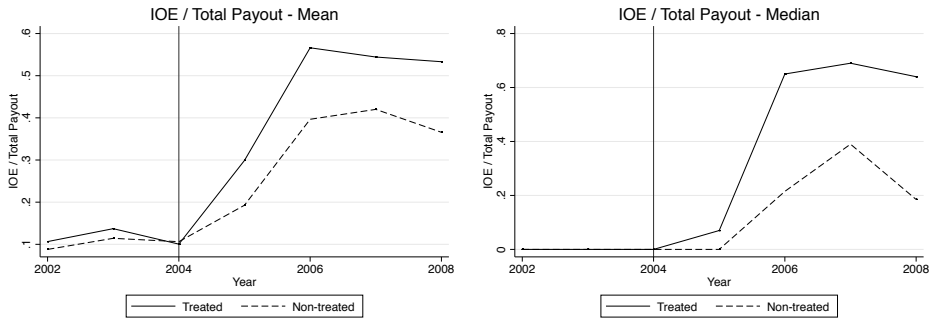
(a) IOE Dummy - Mean

(b) IOE Dummy - Median



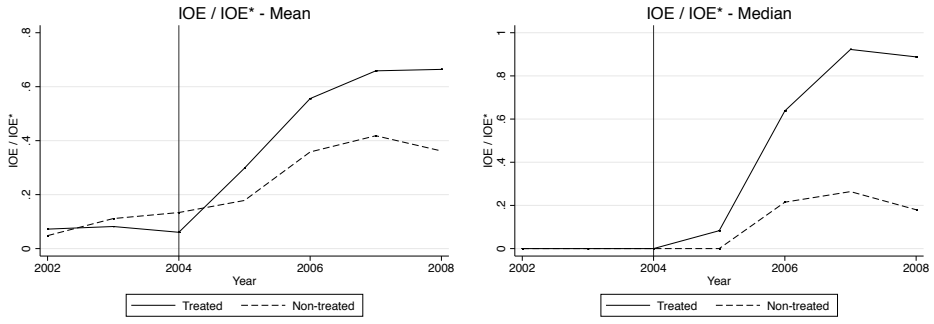
(c) IOE/Total Payout - Mean

(d) IOE/Total Payout - Median



(e) IOE/IOE* - Mean

(f) IOE/IOE* - Median



Source: Authors' elaboration.

Table 5: Results of the Difference in Differences Analysis

Note: This table reports the results of the basic empirical strategy. We divide firms into two groups on the basis of their control structure. The treatment group covers companies with pension funds or investment entities as the first or second largest shareholder with voting rights. The control group consists of similar companies in various dimensions, except for the participation of associative investment entities as the first or second largest shareholder. Time variables are defined by the effectiveness of Law 11,054/2004, the "Pension Funds' Reform" (it took effect in Brazil by January 1, 2005). Therefore, After refers to the period 2005 to 2008 and Before refers to the period 2002 to 2004. In all Panels, we calculate the average of each referred variable pre- and post-law, i.e., we have the averages Before and After the event. Each Panel refers to a different dependent variable, as previously described. The DD in each Panel is the variable of interest. ***, **, and * imply significance at the 99%, 95%, and 90% levels, respectively.

Variables	Before	After	Difference	N
		Panel A: IOE Dummy		
Treat	0.177	0.768	-0.591^{***}	151
Control	0.153	0.534	-0.381^{***}	538
Difference			-0.210^{***}	
		Panel B: IOE / Total Payout		
Treat	0.115	0.485	-0.370^{***}	151
Control	0.103	0.344	-0.240^{***}	538
Difference			-0.129^{**}	
		Panel C: IOE / IOE*		
Treat	0.071	0.541	-0.470^{***}	151
Control	0.099	0.329	-0.230^{***}	538
Difference			-0.240^{***}	

Source: Authors' elaboration.

graphical analysis. Although non-treated firms also started to pay out more using IOE (propensity to pay out using IOE increased from 15.3% to 53.4%), the treated group increased IOE payments even more (propensity increased from 18% to 77%). More importantly, the DD confirms that the difference in IOE payments is both statistically and economically significant. For an average treated firm, the ratio of IOE to total payout (IOE / Total Payout) increased by 12.9% more than that of an average non-treated firm. This result makes sense economically because, before pension funds' reform, only 11.5% of total cash payouts were made through the tax-deductible dividends called IOE. After the event, an average company increased its IOE payments by approximately 321.7%, or 37.0 percentage points. This result implies that, in the after period, almost half (48.5%) of the

cash payments made to firms' shareholders were made through IOE and half through traditional dividends. Similar results (and even statistically stronger) are found related to the ratio between the IOE amount during the year and the maximum allowed by Brazilian fiscal legislation. The average treated firm increased IOE / IOE* by 662.0%, or 37.0 p.p. After Law n. 11,053/2004 came into effect, more than half of the maximum IOE distribution (54.1%) was utilized by treated firms. This sharp increase was also economically and statistically significant in relation to the non-treated group: an average increase in total payments of 24.0 p.p.

Overall, our basic results of firms' response to the new taxation environment are consistent with the tax preference theory of dividend policy. When the marginal tax rate was set to 15% for pension funds — as it still is for individuals, corporations, state-owned firms, and foreign investors — a large proportion of firms controlled by pension funds preferred to use regular dividends instead of IOE. After the tax reform, the 0% tax rate made it significantly more likely for these companies to pay out IOE — more than 80% of treated firms used IOE distributions after 2005. Thus, we argue that taxation at the shareholder level plays a significant role in dividend distributions,[15] a result that supports the hypothesis that taxes are a first-order determinant of payout policy. This result contradicts the survey results of Brav et al. (2005), who found that taxes play a second-order role in firms' payout policy decisions. However, this result is consistent with recent studies on the primary role of taxes on corporate decisions, such as George W. Bush's 2003 U.S. tax reduction (BROWN; LIANG; WEISBENNER, 2007; BRAV et al., 2008), and a surprise increase in corporate taxes for a group of Canadian publicly traded firms (DOIDGE; DYCK, 2015).

1.4.3.3 Evidence from multivariate models

We now continue our analysis by estimating multivariate models and include a set of control variables that could have changed concomitant with the pension fund tax reform and affected the outcome variables. In Table 6, we investigate the impact of Law n. 11,053/2004 on the use of IOE through the standard DD framework. We include industry fixed effects in all regressions to control for industry heterogeneity and year fixed effects to control for aggregate economic shocks.

In column one of Table 6, we report the marginal effects from probit estimations. After the

[15] The neoclassical view of taxes argues that shareholders maximize their after-tax total cash flows, i.e., we should consider both the tax shields at the firm level (marginal tax benefit) and the costs at the investor level (marginal cost of receiving IOE).

Table 6: Multiple regression analysis based on non-linear models (Probit and Tobit)

Note: This table presents the results of the Probit and Tobit panel data regressions for the periods 1998–2004 and 2005–2008, respectively. Dependent variables are IOE Dummy (for Probit regressions) and IOE / Total Payout (for Tobit regressions). Year and industry fixed effects are included in all regressions. Robust and firm-level clustered standard errors are included. The estimated coefficient and the z statistic (in parentheses) are reported for each variable. ***, **, and * imply significance at the 99%, 95%, and 90% levels, respectively.

Variables	IOE Dummy				IOE / Total Payout			
	Probit 1	Probit 2	Probit 3	Probit 4	Tobit 1	Tobit 2	Tobit 3	Tobit 4
After × Treat	0.302***	0.382***	0.392***	0.269**	0.076***	0.091***	0.095***	0.049***
	(2.73)	(3.07)	(3.11)	(2.13)	(2.71)	(3.22)	(3.36)	(2.70)
ln (Total Assets)		0.153***	0.151***	0.283***		0.051***	0.050***	0.130***
		(5.44)	(5.41)	(8.62)		(7.20)	(6.95)	(11.97)
Debt / Total Assets		0.0404	0.0375	0.697**		−0.108*	−0.112*	0.201
		(0.18)	(0.16)	(1.98)		(−1.82)	(−1.88)	(1.63)
EBIT / Total Assets		1.364***	1.337***	1.267*		0.096	0.081	0.390
		(3.18)	(3.13)	(1.80)		(0.69)	(0.58)	(1.52)
Good Governance		−0.118	−0.117	−0.230***		−0.031	−0.031	−0.106***
		(−1.51)	(−1.48)	(−3.60)		(−1.23)	(−1.18)	(−3.05)
% Cash Flow Largest Shareholder		−0.314*	−0.176	−0.374		−0.134***	−0.067	−0.102
		(−1.78)	(−0.69)	(−1.43)		(−2.79)	(−0.71)	(−1.06)
% Excess Cash Flow Largest Shareholder		−0.295	−0.198	0.745		−0.053	0.022	0.306
		(−1.34)	(−0.44)	(0.78)		(−0.95)	(0.16)	(1.07)
(% Cash Flow Largest Shareholder)^2			−0.144	−0.0764			−0.070	−0.064
			(−0.52)	(−0.41)			(−0.76)	(−0.90)
(% Excess Cash Flow Largest Shareholder)^2			−0.248	−0.299			−0.175	−0.188
			(−0.30)	(−0.19)			(−0.67)	(−0.35)
After ×[ln (Total Assets)]				−0.210***				−0.108***
				(−6.56)				(−9.33)
After × (Debt / Total Assets)				−0.804**				−0.345***
				(−2.11)				(−2.70)
After × (EBIT / Total Assets)				−0.211				−0.317
				(−0.27)				(−1.17)
After × Good Governance				0.416*				0.102***
				(1.92)				(2.69)
After × % Cash Flow Largest Shareholder				0.647				0.104
				(1.24)				(0.70)
After × % Excess Cash Flow Largest Shareholder				−1.052				−0.359
				(−1.08)				(−1.21)
After × [(% Cash Flow Largest Shareholder)^2]				−0.325				−0.004
				(−0.68)				(−0.03)
After × [(% Excess Cash Flow Largest Shareholder)^2]				0.333				0.195
				(0.19)				(0.35)
Constant	−1.863***	−7.851***	−7.813***	−22.51***	−0.901***	−2.867***	−2.838***	−10.86***
	(−5.84)	(−7.02)	(−7.03)	(−5.02)	(−4.79)	(−5.52)	(−5.47)	(−5.85)
N	676	592	592	592	682	596	596	596
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Clustered Standard Errors	Firm level	Firm level	Firm level	Firm level	Firm level	Firm level	Firm level	Firm level
Pseudo R-Sq	0.272	0.405	0.406	0.466	0.193	0.271	0.272	0.357

Source: Authors' elaboration.

implementation of the new law, firms with high pension fund participation on voting shares increase their probability of using IOE by an average of 30.2%. In column two, we include a set of control variables that refer to firm characteristics such as size, leverage, profitability, and liquidity. Further,

we control for ownership concentration and separation between control and cash flow. On the basis of theoretical predictions, we also include a quadratic function of both ownership concentration and deviation from control and cash flow. Finally, because our data are not matched in these regressions, we add the interaction between each core control variable and the after-reform variable (year $\geq$ 2005), in column 4. Although the magnitude of the marginal effect of the interaction (treat $\times$ post) variable varies from 30.2% to 41.1%, the results continue to be statistically significant at the 1% level and the qualitatively nature of our results remains unaffected in all probit estimations. Taking Probit 3 model as a parameter, we estimate that an increase from 0 to 1 in the treat $\times$ post variable increases the probability of paying IOE of 41.1%, all else being equal.[16] The magnitude of the change in the probability reveals that the result is not only statistically significant but also economically significant.

Moreover, Table 6 reveals a strong relation between size and the probability of paying IOE. Consistent with the evidence that small firms are less likely to use tax planning instruments, the coefficient estimate in column 3 suggests that a one standard-deviation increase in the logarithm of total assets (1.79) at the mean leads to a $(1.79 \times 0.151 =)27.0\%$ increase in the probability of distributing IOE. We also document a positive and statistically significant impact of profitability on the propensity to pay out IOE. According to the marginal effect estimated in column 3, a one standard deviation increase in EBIT over Total Assets (0.07) leads to a $(0.07 \times 1.337 =)9.4\%$ increase in the probability of using IOE to distribute earnings to shareholders. It is possible that more profitable firms are likely to pay more taxes; therefore, the marginal benefit of paying out IOE instead of regular dividends is higher.

In conclusion, the pension funds' reform increased the probability of treated firms using IOE payments instead of regular dividends. This difference is significant even when controlling for a large number of firm characteristics and industry and year fixed effects. We estimate that, after the reform, treated firms are 41.1% more likely to pay out IOE relative to an average non-treated firm. We also find that size (+) and profitability (+) are economically and statistically significant factors that affect the propensity of a given firm paying out IOE.

Next, in columns 5–8, we repeat the same analysis previously described, but using a continuum variable as a measure of IOE (IOE / Total Payout). We estimate this variable using Tobit

[16] All of these estimations are based on the calculated marginal effects of each variable in the model.

regressions because our dependent variable is censored in zero. From column 5, we observe that the point estimate for the interaction variable (treat $\times$ post) is 0.076. Therefore, the regulatory change increased the ratio of IOE payments to total cash payments in 7.6%, on average. Even including the same control variables as described in columns 1–4, we verify that the point estimates are very close (ranging between 4.9% and 9.5%). Therefore, we estimate that the reform caused a change in the IOE payments of treated firms of approximately 9% of the total cash payments that would have occurred in the absence of the reform.

Finally, because our baseline sample is unmatched,[17] we follow Vig (2013) and proceed in columns 4 and 8 with the interaction between the dummy variable for the post period (post-reform) and the core control variables. In addition to size being a positive and strong determinant of IOE for the entire period, we find that size has a negative effect on the probability of a firm paying out IOE after the reform. We also find a significant difference in the relation between IOE and Debt after the reform — a one-standard deviation in Debt / Total Assets $\times$ Post (0.29) leads to a $(0.29 \times -0.804 =) - 23.3\%$ increase (or 23.3% decrease) in the probability of a given firm paying IOE. This result is consistent with the notion that firms jointly determine its financial policies (JENSEN; SOLBERG; ZORN, 1992; GIVOLY et al., 1992; GATCHEV; PULVINO; TARHAN, 2010; LIN; FLANNERY, 2013). Specifically, we find that a sudden decrease in the marginal costs of paying out IOE causes firms to use simultaneously a lower debt tax shield (Debt / Total Assets) and a higher IOE. We discuss the joint effects and the main implications of DTS and ETS in greater detail in Section 5.

1.4.4 Falsification tests (Placebo periods)

As discussed by Roberts and Whited (2013), because the key assumption behind the DD estimator — the parallel trends assumption — is untestable, some robustness tests should be performed to secure the internal validity of our results. The authors suggest repeating the DD analysis on years before the real event, i.e., to proceed with falsification tests. In these placebo periods, the estimated treatment effect should be statistically indistinguishable from zero to ensure that the observed change is a result of the treatment and not an alternative force. The same intuition was

[17] In Table 4, we show that a large number of observable characteristics are similar for treated and non-treated firms, such as size, profitability, investment opportunities, and so on.

applied to the placebo test performed by Almeida et al. (2012).

We redid our DD analysis using two years before the real event as a falsification test. Therefore, instead of starting on January 1, 2005, we simulate a situation in which Law 11,053 took effect on January 1, 2003 and January 1, 2004, respectively. We calculate the mean of the outcome variables for treated and control firms in the year after and the year before the treatment.[18]

Table 7 shows the results of the placebo tests. Overall, we observe that the placebo results contrast with our main results. In Placebo $\#1$, which simulates the effectiveness of the law on January 1, 2003, the DD estimates are statistically indistinguishable from zero in all variables: IOE Dummy (0.9 p.p.), IOE / Total Payout (0.5 p.p.), and IOE / IOE* (-5.3 p.p.). The same phenomenon occurred in the Placebo $\#2$: -2.5 p.p., -2.9 p.p., and -4.5 p.p., respectively. Therefore, treated and control firms have virtually identical dividend behavior in 2003 and 2004. More broadly, the results of these falsification tests help us rule out alternative explanations for the results reported in Table 7.[19]

1.5 Do equity tax shields substitute debt tax shields?

We continue our analysis by assessing the relation between DTS and ETS — Interest on Equity (IOE). Because IOE payments are interest expenses, similar to the interest on debt payments, one could expect a substitution effect between debt and IOE for tax reasons, especially after Law 11,053/2004 was implemented and firms started to increase IOE payments. Intuitively, firms could keep the same tax benefits from interest payments by simultaneously increasing IOE payments and reducing leverage. Our hypothesis is that firms jointly determine their financial policies, i.e., one exogenous tax variation that directly affected dividend policy can indirectly affect debt financing through the interdependence of corporate financial policies.

[18] Because our treated and control definition is based on ex-ante ownership structure information (pre-treatment period), we keep the same firm characteristics to define treated and control units, i.e., firms in both groups are similar in dimensions such as size, profitability, leverage, liquidity, and ownership structure, but control firms do not have pension funds as a first or second largest shareholder.

[19] One could argue that unobservable characteristics could cause both pension funds' ownership and the choice between regular dividends and IOE. If so, we should observe the same behavior (treated firms increase more IOE payments) in the years before the real event. However, our results of the placebo tests suggest that this is not the case.

Table 7: Results of the falsification tests (Placebos)

Note: This table reports the results of the Placebo Tests using the standard DID approach (same as used in Table 5). The DD refers to the variable mentioned in each panel. ***, **, and * imply significance at the 99%, 95%, and 90% levels, respectively.

Variables	Before	After	Difference	Before	After	Difference
			Panel A: IOE Dummy			
	Placebo #1: Event Year 2003			Placebo #2: Event Year 2004		
Treat	0.150	0.190	0.040	0.190	0.190	0.000
Control	0.123	0.154	0.031	0.154	0.179	0.025
Difference			0.009			−0.025
			Panel B: IOE / Total Payout			
	Placebo #1: Event Year 2003			Placebo #2: Event Year 2004		
Treat	0.106	0.137	0.031	0.137	0.100	−0.037
Control	0.088	0.114	0.026	0.114	0.106	−0.008
Difference			0.005			−0.029
			Panel C: IOE / IOE*			
	Placebo #1: Event Year 2003			Placebo #2: Event Year 2004		
Treat	0.072	0.082	0.010	0.082	0.060	−0.022
Control	0.048	0.111	0.063	0.111	0.134	0.023
Difference			−0.053			−0.045

Source: Authors' elaboration.

1.5.1 Empirical evidence

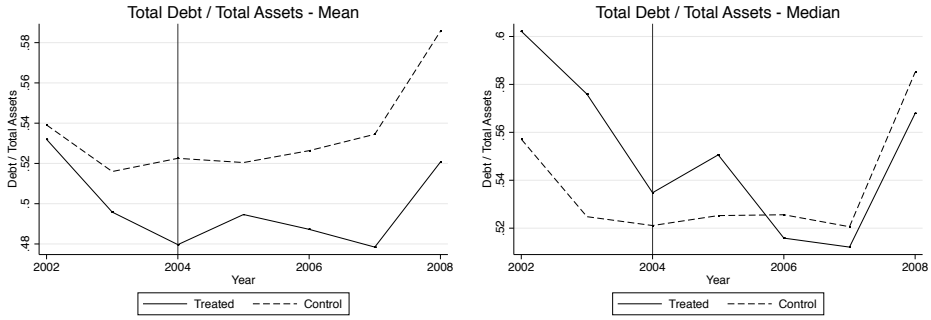
We estimate from Table 6 that a one-standard deviation in Debt / Total Assets $\times$ Post leads to a -23.3% increase (or 23.3% decrease) in the probability of a given firm paying IOE, all else being equal. As is noted, Law 11,053/2004 did not directly affect firms' debt incentives. However, by reducing the marginal costs of IOE payments for pension funds, firms with large pension fund shareholders (treated firms) became approximately 39% more likely to use IOE than the average control firm. This increasing use of ETS could lead to a decrease in the use of DTS through substitution between DTS and ETS. In this case, the reform could have an indirect effect on invested firms' leverage decisions.

Our graphical evidence reported in Figure 4 suggests that, after the reform, a substitution

Figure 4: Mean and median of Debt / Total Assets and ln (Interest Expenses) for treated and non-treated firms, 2002–2008

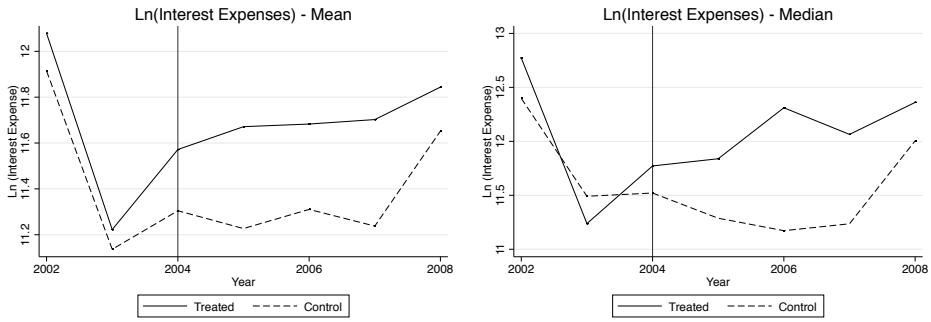
(a) Total Debt / Total Assets - Mean

(b) Total Debt / Total Assets - Median



(c) ln (Interest Expenses) - Mean

(d) ln (Interest Expenses) - Median



Source: Authors' elaboration.

effect indeed exists between IOE payments and debt. Treated and control firms are determined using the same criteria as before: the treated group is formed by companies with pension funds or investment entities as the first or second largest shareholder with voting rights, and the control group consists of similar companies in various dimensions, except for the participation of associative investment entities as the first or second largest shareholder. We report the mean (left side) and median (right side) of each group using two continuous variables: Total Debt / Total Assets (upper graphs of the figure) and the logarithm of total interest expenses (which includes debt and non-debt interest payments).

Consistent with the substitution hypothesis, we observe in Figure 4 that the treated group slightly reduced the Debt to Assets ratio (0.498 to 0.486, on average) and the control firms increased the Debt to Assets ratio (0.525 to 0.539). If we consider the median instead of the mean, the differences are even higher: 0.582 to 0.534 in the treated group and 0.537 to 0.527 in the control

group. On average, treated firms reduced Debt / Total Assets by $([0.486-0.498]-[0.539-0.525] =)2.6$ p.p. relative to control firms. The average Debt / Total Assets before the treatment period implies an economically significant reduction in Debt / Total Assets of $(0.026/0.498) = 5.22\%$ for treated firms.[20]

In addition to the fact that treated firms reduced their leverage relative to control firms after the reform, the graphics at the bottom of Figure 4 suggest that treated firms increased their total interest expenses relative to control firms. Because IOE is reported on firms' financial statements as interest expenses in the same manner as debt interest payments, we interpret that the increase in ETS was greater than the relative reduction in DTS. In other words, treated firms increased their total interest expenses by paying more IOE to its shareholders, thus reducing earnings before corporate taxes and paying less corporate taxes. This result is consistent with the substitution between the DTS and ETS hypothesis and suggests that treated firms experienced growth in the use of tax planning instruments after the reform.

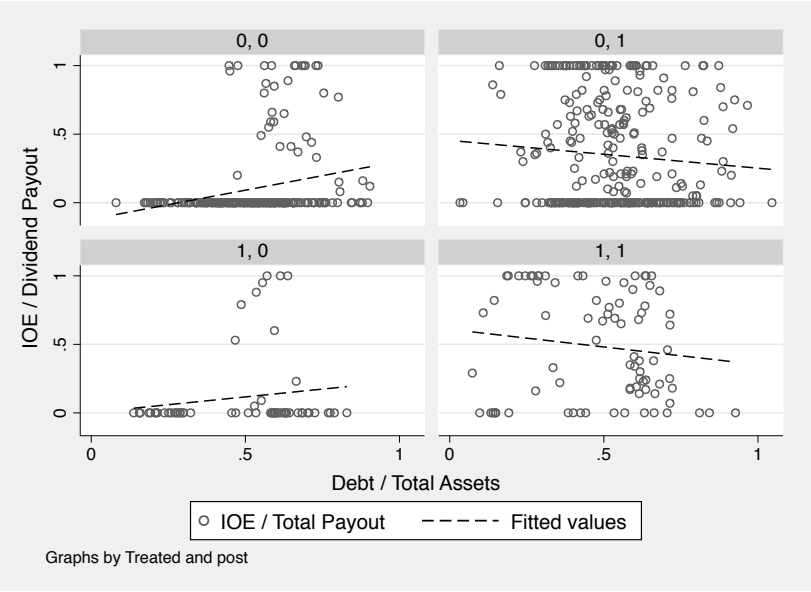
Finally, we report on the scatter plot of IOE / Total Payout (vertical axis) and Debt / Total Assets (horizontal axis) in Figure 5 for treated and non-treated firms and pre- and post-periods. In the first quadrant, we restrict observations to non-treated firms in the post period (0,1). In the second quadrant and following the order, we restrict observations to treated firms in the post-period (1,1), treated firms in the pre-period (1,0), and non-treated firms in the pre-period (0,0).

We observe that, after the reform (post $= 1$, right side graphics), the relation between leverage and the ratio of IOE payments to Total Payout changed significantly. Before the reform (post $= 0$), a positive relationship existed for both treated and control firms, suggesting that these mechanisms were not substitutes. However, after the treatment, this relation became negative, suggesting that firms started to pay more IOE and reduced their leverage, which is consistent with the lower marginal costs associated with IOE relative to debt.

We also run regressions using Debt / Total Assets as a dependent variable (unreported results). In addition to the substitution between IOE and debt, we find that size is positively related and statistically significant to leverage. We also find that leverage is negatively related to profitabil-

[20] Although the context is different, the magnitude of our results is similar to that reported by Vig (2013). The author estimates that treated firms reduced their Secured Debt / Assets and Total Debt / Assets by approximately 5.1% and 4.4%, respectively, after a securitization reform in India.

Figure 5: Relation between IOE / Total Payout and Debt / Total Assets for treated and non-treated firms, pre- and post-periods, 2002–2008



Source: Authors' elaboration.

ity, consistent with other findings in the empirical capital structure literature (see, for example, RAJAN; ZINGALES, 2005; FRANK; GOYAL, 2009). Profitable firms seem to use more internal sources of funds (i.e., retained earnings) to finance their investment needs rather than issue new debt. We also document that liquidity is statistically significant and negatively related to leverage. Liquidity can proxy for investment opportunities and suggests that firms prefer to use internal resources rather than external financing.

1.5.2 Discussion

Although the interdependence of firms' financial decisions is not a new idea in the literature, scarce empirical research exists on this topic. Jensen, Solberg and Zorn (1992) are one of the first authors to address the direct and indirect relation among corporate financial policies. More recently, Gatchev, Pulvino and Tarhan (2010) make theoretical arguments and find empirical evidence that ignoring the interdependent nature of financial decisions results in misleading and often incorrect results. Consistent with our evidence in this paper, Kolay, Schallheim and Wells (2011) find empirical evidence that firms use non-DTS to reduce taxable income. Our results also relate

to the literature that studies firms' reactions to changes in taxation at the shareholder level. Lin and Flannery (2013) estimate that the 2003 U.S. tax reduction for individuals caused affected firms' leverage to decrease by approximately 5 percentage points. Although in a different context, our results are very similar to those of Lin and Flannery (2013) in both direction and magnitude.

Overall, our empirical evidence is consistent with the hypothesis that firms jointly determine their financial policies (JENSEN; SOLBERG; ZORN, 1992; GIVOLY et al., 1992; GATCHEV; PULVINO; TARHAN, 2010; LIN; FLANNERY, 2013). In addition, our shock-based empirical strategy allows us to infer causality between a taxation change at the investor level and firm responses to both dividend policy and capital structure. The magnitude of our results also suggests that firms' shift from DTS to non-DTS is both statistically and economically significant — the average treated firm increases the probability of paying IOE by 39.2% and reduces Debt/Assets by 5.2%. In this sense, our paper also contributes to the literature by providing new evidence on the interdependence of corporate financial decisions, which may amplify the effects of corporate reforms on the entire economy.

1.6 Conclusions

Does a change in taxation at the investor level cause changes in firm-level outcomes? The results of this paper suggest that the answer is yes. We explore a unique setting in Brazil to address this question. More specifically, we analyze a pension fund reform in Brazil that changed the marginal income tax rate on the receipt of interest on equity (IOE) to pension funds, all else being equal for other shareholders. We find evidence that invested firms with equity participation of pension funds as the first or second largest shareholders began to distribute a larger share of their earnings in the form of IOE compared with a counterfactual of similar firms.

Whereas the survey results from Brav et al. (2005) show that taxation plays a secondary role in corporate finance, our evidence supports the hypothesis that taxes actually have a first-order impact on a firm's financial decisions. We also highlight the channel through which this impact occurs: high ownership concentration allows controlling shareholders to enforce changes in firms' decisions through their close ties to either executives or members of the supervision board.[21]

[21] The recent paper written by Foz, Kim and Kronlund (2014) find that insider ownership affects firms' responses to changes in the tax environment.

However, we emphasize that we do not realize a welfare analysis that could document the net effects of the changes observed in a firm's dividend policy. According to the Brazilian tax code, individuals, associative investment entities, and foreign investors reduced their total taxes and, therefore, were better off. However, in certain situations, corporate shareholders may suffer from higher taxation if the invested firm switches its cash payouts from dividends to IOE.[22] Therefore, the net effects for shareholders are very difficult to measure.

We found that the changes in the composition of cash payout distributions are statistically and economically significant. After the reform, treated firms were 39.2% more likely to distribute cash as IOE instead of as regular, non-deductible dividends. These firms also increased the share of IOE in Total Payout by 12.9 p.p. relative to the control group, which represents an 112.2% increase over the pre-treatment average. We also found evidence that treated firms reduced their Debt / Assets ratio by 2.6 p.p. (5.2%) after the reform, consistent with a substitution between debt and IOE for tax purposes. These results indicate that firms jointly determine their dividend and debt policies, i.e., corporate finance decisions are interdependent. From the point of view of policy makers, our analysis provides evidence that tax reforms can generate broader effects on the economy through the interdependence between corporate actions.

Finally, our results provide a better understanding of the reasons why a number of Brazilian companies prefer to pay out their earnings as regular dividends instead of tax-deductible IOE. We show that tax considerations at the investor level are key determinants of a firm's choice between IOE and dividends. Henceforth, one plausible explanation for the number of Brazilian companies that do not use IOE is that their controlling shareholders individually do not have incentive to do so, i.e., the costs they face at the investor level can be higher than the tax benefits of IOE at the invested firm. In this sense, our results also shed light on tax issues that may render ineffective Allowance for Corporate Equity systems, such as the one used in Brazil.

[22] This situation occurs when corporate shareholders cannot compensate for the tax expenses related to the IOE received using firms' annual corporate taxes. The situation also occurs when a corporate shareholder has no positive net income in the same fiscal year in which she receives IOE or when the firm has accumulated losses from previous years that can generate fiscal gains that are carried forward in the same fiscal year.

2 Do Foreign Portfolio Capital Flows Affect Domestic Investment? Evidence from Brazil

Abstract

Although there are several direct and indirect theoretical channels through which financial integration may affect domestic investment, empirical evidence is relatively scarce and still inconclusive. In this paper, we estimate the Brazilian real gross capital formation on a monthly basis and use a VARX framework to assess the impact of Equity Foreign Portfolio Investment (EFPI) on domestic investment growth. Despite the high persistence of the aggregate investment in Brazil, our results suggest that EFPI played a non-negligible role in explaining aggregate investment fluctuations, but only before the 2008 global financial crisis. After this crisis, a period that coincides with an increasing government intervention in the Brazilian economy, unexpected shocks to EFPI led no real effects on investment growth. Our evidence also suggests that the causality direction between the variables is likely to be from EFPI to investment, and not the other way around. Finally, we discuss the importance of EFPI in financing expansions of capital stock in the economy, particularly future infrastructure projects in light of the new Brazilian fiscal reality.

Keywords: Foreign Portfolio Capital Flows, Financial integration, Aggregate investment.
JEL Codes: C53, E43, G17.

2.1 Introduction

Developing countries have historically struggled with insufficient supply of affordable capital to finance investments. A promising agenda to tackle the problem was advocated by the IMF and The World Bank in the late 1980ies, within the umbrella of the so-called Washington Consensus. Developing countries were advised to implement capital account liberalizations, allowing foreign equity capital to flow in, thus promoting integration with global equity markets and financing new capital stock with foreign funds. However, as noted by Aizenman, Pinto and Radziwill (2007), such recipes for growth eventually became the single most controversial policy prescription, and concerning fostering investments, for most of emerging markets, there is no evidence of a growth bonus associated with increasing the financing share which is done with foreign funds. In this paper, we contribute to this spicy debate by investigating whether foreign portfolio capital flows stimulate investment, using Brazil as a case study.

Increased levels of financial integration propelled by foreign equity capital flows allegedly reduce the cost of equity capital in developing markets for the interplay between four main factors: improved risk sharing among local and foreign investors, alleviation of financial constraints as more foreign capital becomes available, increased stock market liquidity and adoption of better corporate governance practices by local firms to attract more sophisticated foreign shareholders. As emerging countries move from financial autarky to become integrated, physical investment should increase accordingly, as a lower cost of equity capital brought about by financial integration expands the portfolio of positive NPV investments in the economy, for cash flows from new investments are now discounted under a lower pricing factor (LEVINE; ZERVOS, 1998; HENRY, 2000; CHARI; HENRY, 2004; BEKAERT; HARVEY; LUNDBLAD, 2005; STULZ, 2005).

Theories justifying increased investment under integration are reasonably sound, but in practice the story is more complicated. Instead, foreign portfolio capital is often blamed for disrupting local financial markets, for its short-termed nature exacerbates volatility and instability, actually hindering new investment because firms are reluctant in expanding their capital stocks when they do not trust foreign capital will stay long (STIGLITZ, 2000; SINGH; WEISSE, 1998). In fact, recent empirical evidence shows that during periods of financial instability, like in the 2008 global financial crisis, foreign equity investors reallocated massive quantities of portfolio capitals

from emerging economies to advanced economies (FRATZSCHER, 2012). As adjusting capital stocks is costly, uncertainty about equity valuations caused by foreign capital sudden reversals might actually discourage new investment. Also, portfolio investment may harm the economy in emerging markets for its pro-cyclicality, as it increases when economies are booming but rapidly retreats when economies are slowing, for overheating exchange rates and for inducing bubbles in real estate and financial asset prices (AIZENMAN; PASRICHA, 2013).

A few papers have studied the effects of financial integration on investment, yet most of these studies built on sets of liberalization events (HENRY, 2000; LAEVEN, 2002; BEKAERT; HARVEY; LUNDBLAD, 2005; CHARI; HENRY, 2008). An important assumption underlying the argument that liberalizations boost investment is that integration occurs instantaneously, or that risk sharing takes place rapidly and the cost of capital declines right after markets are liberalized. In general, these papers report a positive effect of integration on investment and growth, but in most cases these effects refer to short periods after liberalizations. Alternatively, another stream of research views financial integration as a time-varying process, asserting that it takes time for risk sharing to kick in. Domestic market segmentation decreases over time and not overnight, thus reductions in the cost of equity capital are not instantaneous, kicking in progressively as markets further integrate (DE JONG; DE ROON, 2005; CARRIERI; ERRUNZA; HOGAN, 2007; BUCKBERG, 1995). Our paper differs from liberalization studies, as we take a relatively longer-term approach, covering twenty years of foreign equity capitals continuously flowing to the Brazilian equity market, from 1996 till 2015.

The Brazilian experience is an interesting story to study. Like many emerging markets, Brazil experienced a surge in foreign capital flows in the 1990s (CARDOSO; GOLDFAJN, 1998). More recently, in years 2009/2010, increases in capital flows raised concerns related to financial stability and exchange rate overheating, to which Brazil responded with several capital controls on equity and fixed income investments (CHAMON; GARCIA, 2016; JINJARAK; NOY; ZHENG, 2013). Moreover, Brazilian private firms long suffer from credit constraints, relying heavily on internally generated cash flows to finance investments (TERRA, 2003). As a response to the 2008 financial crisis, the Brazilian government has sharply increased the supply of subsidized credit from state-owned banks (mainly through BNDES) to the private sector, especially for large firms, a policy of betting on so-called national champions firms, hoping to give a boost to investment.

Nevertheless, as of the present moment, there is no evidence this policy has produced any stimulus on investment, but it has notably contributed to the deterioration of fiscal deficits.

As the Brazilian economy faces the toughest recession of its modern history by the time we write, the country urgently needs alternatives to propel investment and deliver growth to make its way out of this crisis. In theory, foreign portfolio capitals offer a promising channel to finance expansions of private capital stock. In fact, there is evidence showing that foreign capitals increase equity valuations and decrease the cost of capital in Brazil (TABAK, 2003; REIS; MEURER; SILVA, 2010; SANVICENTE, 2014; LONCAN; CALDEIRA, 2015), giving a first indication that foreign equity financing might offer a viable solution to overcome depressed investment. However, whether foreign capital has truly contributed to financing new investments remains to be investigated. In our paper, we address this important question.

Our first step in modeling the relation between foreign capital and aggregate investment is to construct a monthly estimate of the Brazilian quarterly gross capital formation series. We do that for two main reasons. First, because we have monthly available information on the main components of the quarterly aggregate investment, following the most recent guidelines of the *System of National Accounts 2008* (UNITED NATIONS, 2009). Second, all other macroeconomic variables we use in the models are available on a monthly basis, including the foreign investment flow. In doing this monthly interpolation of the quarterly investment series, we do not change the properties of the original series and we significantly increase the number of degrees of freedom in our models, which allows us to enhance the number of estimated parameters and thus improve the fit of the model.

Following this procedure, we model the effect of foreign portfolio equity capitals on real domestic investment by developing a monthly vector autoregressive model with exogenous variables (VARX). We follow economic theory and standard neoclassical models of investment (see, e.g., ROMER, 2012) and consider investment, foreign capital, stock market valuation, real interest rates, the exchange rate between the Brazilian Real and the U.S dollar and the supply of credit as a share of gross domestic product as endogenous variables, modeling them simultaneously according to the transmission mechanisms as follows. Neoclassical models of investment predict investment to be an increasing function of future expected profits, which are embedded in equity valuations.

As foreign equity capital decreases the cost of capital, equity valuations soar, expanding the investment opportunity set in the economy. Such higher equity valuations attract new foreign equity investment because foreign investors foresee good opportunities to reap capital gains as stocks appreciate. We also include a potentially negative effect of real interest rates and real exchange rates on investment that are considered endogenous in the model. Finally, we add a vector of exogenous variables in our analysis to capture terms of trade shocks and the global business cycle, such as the IMF commodity prices index, MSCI global stock returns, and a dummy variable capturing recession periods faced by the Brazilian economy.

Our identification strategy on the VARX estimation relies on economic grounds. Because aggregate investment obeys physical constraints, a firm may decide to invest, but measured investment responds with a delay because of the investments' inherent life cycle (KILLIAN, 2011). Therefore, in our recursive ordering of the endogenous variables in the system – which reduced-form errors we orthogonalize applying the Cholesky decomposition on the residual variance-covariance matrix –, we consider aggregate investment as the top variable in the VAR system.[23] Because variable ordering is an important aspect in recursively identified models such as ours', we follow Killian (2011) and set our recursive structure based on economic justification, specifically, tracking the predictions of the neoclassical model of investment (ROMER, 2012).

The findings from our empirical analysis show that equity foreign portfolio investment (EFPI) tends to produce beneficial effects on domestic investment. When analyzing the full period covered in our study (1996-2015), we find that a 1 percentage point (p.p.) increase in EFPI to GDP ratio rose investment (gross capital formation) growth by 0.3% fifteen months ahead, keeping all else equal. This effect, however, changes dramatically when we split our sample in before and after the 2008 financial crisis. Before the crisis, the response of aggregate investment growth to a 1 p.p. increase in EFPI/GDP was positive and significant (about 0.4%). In the after crisis period, foreign capital no longer boosted domestic investment - the cumulative impulse-response functions (COIRFs) show that a that a one percentage point increase in EFPI/GDP decreased aggregate investment growth in about 0.3%.

Although we can not state this specific channel, this duality between the two analyzed time

[23] In our recursive VAR model, it implies that investment can not be contemporaneously affected by a shock on other endogenous variables. Investment can, however, respond to lagged innovations on the other endogenous variables.

windows is consistent with the idea that excessive government intervention in credit markets can neutralize the impact of EFPI on investment, as transfers from Brazilian National Treasury to State-owned Banks increased from 0.9% of GDP in 2008m8 to 9.8% of GDP in 2015m10. Meanwhile, unconventional monetary policies in developed countries (low interest rates since the beginning of the crisis) may have contributed to equity flows to remain at a relatively high level in Brazil, despite growing public fiscal deficits, lower business confidence and deteriorating economic fundamentals. Taken together, these two explanations may help to explain why the role of EFPI on aggregate investment changed significantly after the 2008 financial crisis.[24]

Furthermore, the forecast error variance decomposition (FEVD) analysis shows that only 3% of movements on aggregate investment in a horizon of fifteen months can be explained away by foreign capital flows, whereas 84% of such variations are due to lagged investment. Given foreign investments account for nearly 25% of total stock market capitalization, the contribution of foreign capital in financing physical investments, though positive and relevant, seems inefficiently low in economic terms. We observe a similar result even before the crisis, when EFPI had a more pronounced effect in aggregate investment: from 2008m12 to 2015m10, only about 4% of the error in the forecast of Brazilian gross capital formation growth is attributed to EFPI. These results suggest that, albeit statistically significant, the impact of EFPI on gross capital formation is economically modest.

To deal with a potential feedback effect between investment and EFPI, we proceed to a Granger causality analysis. Our results suggest that the direction of causation occurs from EFPI to investment, and not the other way around (i.e., we find a unidirectional effect from EFPI to investment). However, just as suggested by the COIRFs and FEVDs analyses, lags of EFPI are only useful for predicting aggregate investment before the 2008 crisis. After the crisis, we can not reject the null hypothesis that lagged values of EFPI are jointly equal to zero in the model.

Finally, we perform a sensitivity analysis of our results. Following Sims (1981) and Lütkepohl (2011), we check if our evidence is robust to different variable ordering in the Choleski decomposition of our VARX system. Assuming aggregate investment at the top and EFPI at the bottom[25]

[24] This parameter changes we find after the crisis is also consistent with recent empirical evidence in VAR models and evaluation of shock transmission (AASTVEIT et al., 2016).

[25] As discussed later with more details, we assume that aggregate investment can not react contemporaneously to shocks in other endogenous variables because of its inherent rigidity, i.e., it takes time to be executed. On the other

of the vector of K=6 endogenous variables, we test all the 4! = 4 x 3 x 2 x 1 = 24 different model specifications and find that the cumulative responses of aggregate investment to a shock in EFPI is positive in 100% (24/24) of the combinations. Moreover, consistent with our previously evidence, the effect of EFPI in investment is more pronounced in the before-crisis period.

In light of this empirical findings and of existing theories of corporate ownership structure and financial contracting, we discuss how interactions between corporate and public policies may help firms in making the most of the foreign equity investments they receive. We also debate the role of external equity capital in financing investments in infrastructure, in light of the restrictive fiscal reality faced by Brazil in the current period and very likely in years to come.

As side contributions, we also find evidence that increased stock market valuations attract more foreign capital flows, what we interpret as an additional component to virtuous cycle between stock market returns, foreign equities, and investment: high equity market valuation induce foreign portfolio capitals, which in turn stimulate investment. There is a side effect, though. Foreign funds cause the exchange rate to appreciate, potentially undermining the competitiveness of the exports sector, for example. But naturally the reverse of such a currency side effect is that it also becomes cheaper to import capital goods, so much needed to modernize the existing capital stock and improve productivity in Brazil, a country which has been on a 50 years productivity snooze according to *The Economist* magazine. Overall, the results from our study corroborate the argument that foreign capitals might be helpful in funding investments, leaving an interesting message for economic policy makers, and adding a new piece of evidence to this long academic debate on the pros and cons of capital account liberalizations.

The rest of our paper is organized as follows. Section 2 brings our model specification, data and econometric strategies employed. In section 3 we describe our results, and in Section 4 we discuss its main implications in the Brazilian context. Finally, in section 5, we present the conclusions of our paper.

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hand, because of its flexibility (foreigners transactions in securities are typically very liquid), we assume that EFPI can react immediately to shocks in other variables, such as investment, credit to GDP, real interest rate, real exchange rate, and Ibovespa.

2.2 Methodology and Data

In this subsection, we present our methodology and a description of our data. We focus on the estimation process of a monthly investment series for the Brazilian economy and the modeling of the relationship between the EFPI and the aggregate investment.

2.2.1 Estimating the Brazilian monthly aggregate investment series

The aggregate investment rate of the Brazilian economy is released by the Brazilian Institute of Geography and Statistics (IBGE) on a quarterly basis, following the publishing schedule of the country's System of Quarterly National Accounts (SQNA). Investment is defined as the gross fixed capital formation and consists of outlays on additions to the fixed assets of the economy plus net changes in the level of inventories (UNITED NATIONS, 2009).

Although investment is only available on a quarterly basis, the IBGE itself monthly discloses the industrial production level of Brazilian industry sector according to the goods' category of use, comprising equipment (gross capital formation), intermediary goods (intermediate consumption), and consumer goods (durable, non-durable and semi-durable). If the real aggregate investment is not observable on a monthly basis, production of capital goods - which is a proxy for the increase in gross fixed capital formation - is available monthly through the Monthly Industrial Survey (PIM/IBGE). We also observe on a monthly basis another important component of the domestic aggregate investment: the production of standard construction inputs, such as cement, iron, steel, among others, through another table of the Monthly Industrial Survey (PIM/IBGE). According to the IBGE, from 2010 to 2014, capital goods and construction industries' accounted together for about 89% of gross total capital of the country.[26]

We then construct a monthly investment series based on the evolution of both capital goods and construction inputs production, aggregating the referred series accordingly to its weight in the 2010-2014 average gross value added. Because these monthly available series represents the evolution of almost 90% of the gross fixed capital formation, one could expect both monthly and quarterly series to be highly correlated.[27] However, because both series does not fit perfectly, we

[26] The other 11% refers to intellectual property products (IPPs - almost 11%) and net changes in the level of inventories (less than 0.5%).

[27] Indeed, our high-frequency monthly estimate of investment is highly correlated with the original quarterly gross

improve the our monthly estimation by applying the Denton's proportional method (DENTON, 1971) to interpolate the quarterly investment series with the high-frequency investment indicator. This method is described as "[...] relatively simple, robust, and well suited for large-scale applications." (BLOEM; DIPPELSMAN; MAEHLE, 2001, p. 98). It is important to note that the indicator series only contribute to the pattern of the interpolated points, not modifying the characteristics of the original series (the quarterly aggregate investment of the Brazilian economy). Technically, the method is a constrained least squares problem, in which the interpolated series obeys the original low-frequency totals (which represents the imposed constraint). Because of these advantages, Denton's proportional method is widely used in countries' SNAs around the world.

The results of the application of the Denton's proportional method are shown in Figure 6. While the line plot represents the estimated monthly series, the scatter plot refers to the original quarterly investment series. Just as expected, both series present the same cyclical pattern and time trends. From 1993m3 to 2005m3, Brazilian investment did not increase its level over time. Starting in 2005, investment rose significantly, stopping its positive slope temporarily during the 2008 financial crisis, but recovering fast and keeping its upward trend until mid-2013. On early 2014, a severe economic recession imposed a downward trend to the aggregate investment, which shows no sign of recovery until our last observation, referred to 2015m10.

The original series in the investment rate is provided by the IBGE from the first quarter of 1996 (1996Q1) to the third quarter of 2015 (2015Q3), totaling 79 observations. We have 238 observations of our monthly series, starting in 1996M1 and ending in 2015M10. For convenience, our series are shown in index numbers, and 2011Q1 is set to be equal to 100.

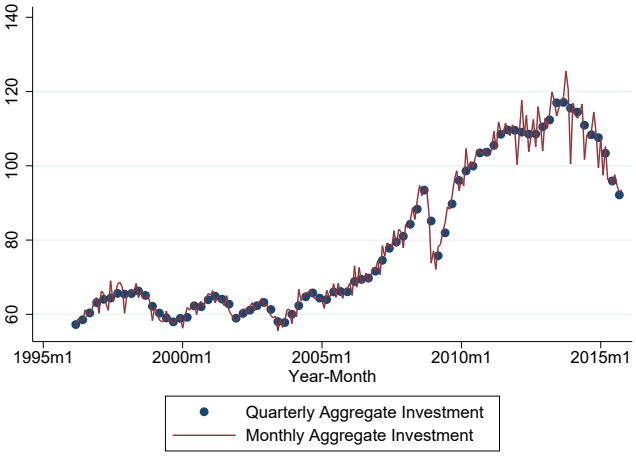
We can infer from Figure 6 that the data is well suited for the investment data. For example, during the 2008-2009 financial and economic crisis, the industrial production of capital goods suffered a sharp decrease, the same pattern observed in the gross formation of fixed capital. One could also note that the volatility of the Brazilian industry production has increased recently. In fact, the economic recession that began in 2014[28] caused a sharp drop in the industrial production, especially in the capital goods sector.

capital formation: the Pearson's linear correlation coefficient during the 1996-2015 period is 0.81.

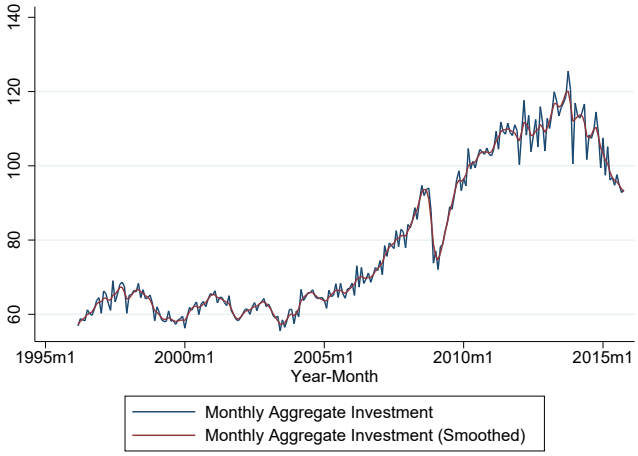
[28] In August 4th, 2015 meeting, the Brazilian Business Cycle Dating Committee (CODACE) identified a local maximum point (peak) in the Brazilian business cycle in 2014Q2, suggesting that the start of the recession occurred in that quarter.

Figure 6: Brazilian Gross Fixed Capital Formation (quarterly) and our high-frequency investment estimate (monthly), 1996 to 2015

(a) Monthly vs Quarterly Investment



(b) Monthly vs Smoothed Monthly Investment



Source: Authors' elaboration.

Note: (a) We compare the original seasonal adjusted quarterly gross fixed capital formation of the Brazilian GDP (blue scatter) with the monthly gross fixed capital formation estimated with the Denton Proportional Method (red solid line), using a high-frequency series composed by production of construction inputs and production of capital goods.

(b)We use the Hodrick-Prescott high-pass filter to separate the monthly investment series into trend and cycle components, with the smoothing parameter $\lambda = 1$. We then use the trend component of the series (red solid line) as a smoothed version of the original monthly investment series (solid blue line).

2.2.2 Data and descriptive statistics

Following the definition of the aggregate investment series, we calculate our measures of net foreign portfolio investment (EFPI) in Brazil using data from the Securities Exchange Commission of Brazil (CVM) and Central Bank of Brazil (BCB). The net (inflow minus outflow) foreign capital is converted from USD to Brazilian Real (R$) using the monthly average exchange rate, and then normalized by the 12 month accumulated GDP.[29] It is important to note that a recent methodological change on National Accounts took place in Brazil, following the IMF recommendations established on the sixth edition of the Balance of Payments and International Investment Position Manual (BPM6). Therefore, our series are the most recent, best available information on foreign capital flows.

In order to estimate the partial effect of the foreign portfolio investment on the domestic investment, we collect data for market characteristics such as local interest rates (real interest rates - RIR, and the annualized spread on the inter-bank deposit rate - SWAP360), real exchange rate (RER), country specific risk (EMBI+ Brazil), investment opportunities (Consumer Confidence Index - CCI, Brazilian stock market representative index - IBOVESPA, and MSCI+ index for both the world and emerging equity markets), U.S. interest rates (Fed Funds Rate and U.S. 10 year bond yield), financial development (Credit-GDP ratio), price of exports and imports (IMF Commodities Price Index and Terms of Trade) and government subsidized credit (mostly transferences from the National Treasury to the Brazilian Development Bank - BNDES). We also include in our models a dummy variable for the recession periods in the Brazilian economy (RECESSION) and a dummy for the 2008 financial crisis. Following the chronology of the global financial recession, this variable takes the value of one starting on September 2008, when Lehman Brothers filing for Chapter 11 bankruptcy protection triggered a spike in interest rate spreads and risk aversion around the world.[30] We therefore set the 2008CRISIS dummy to be equal to 1 if in the period from 2008m09 to 2008m11, and zero otherwise. The summary of the variables we consider in this study is exposed

[29] We collect the GDP accumulated in the last 12 months - current prices (R$ million) from the Time Series Management System of the BCB (series code n. 4382). Because both the numerator (EFPI) and the denominator (12 month GDP) are in current prices, we do not need to adjust this two series by inflation.

[30] The spread between the T-Bill (3m) and the Libor (3m) rose from around 2.8 percentage points in early September 2008 to a peak of almost 5 percentage points in middle October 2008. In November 2008, following the announcement of the Troubled Asset Relief Program (TARP), the interest rate spreads started to decrease to the pre-Lehman Brothers episode levels.

in Table 8.

Table 8: Description of variables

Note: This Table summarizes the variables we use in this paper, from 1996m3 to 2015m10. We classify each variable in groups according to its economic rationale. In column "Label", "Δ" means that the variable is measured in differences of the Natural Logarithm transformed data, and thus is expressed in percentage terms (i.e., Δy_t = [ln(y_t) - ln(y_{t-1})]). If "Δ" is not referred, than the variable is expressed in levels. We follow the results of unit root tests to decide when log-differences are needed.

Variable	Label	Detail	Interpretation	Variable Group	Source
Domestic Investment	ΔINVESTMENT	Monthly estimate of Brazil's real gross capital formation	Domestic Investment	Dependent Variable	Authors' calculation based on IBGE
Net Foreign Porfolio Capital flows	EFPI-GDP	Monthly net foreign capital flows (inflows minus outflows) as a percentage of total GDP	Measure of foreign investor's activity on the stock market	Foreign Capital Flows	CVM and BCB
Real Interest Rate	ΔRIR	Annualized nominal interest rate deflated by the twelve month accumulated inflation	Proxy for Cost of Capital	Cost of financing	BCB
Swap pre-DI	ΔSWAP360	Annualized spread on the interbank deposit rate	Proxy for Brazilian Risk Free Rate	Cost of financing	BCB
Real Exchange Rate	ΔRER	Oficial Exchange Rate adjusted for internal and external price differences over time	Real Price of Domestic Currency (Brazilian Real)	Exchange Rate	World Bank and BCB
Financial Development	ΔCRED-GDP	Supply of credit / GDP	Proxy for financial development and supply of funds to firms	Financial Development	BCB
Ibovespa Monthly Index	ΔIBOV	Local Stock Market Portfolio Index (Ibovespa)	Reflects current and future economic conditions	Investment Opportunities	BCB
EMBI+ BR	ΔEMBI-BR	Spread of Brazilian Bonds over US Bonds (fixed maturity)	Reflects local economy uncertainty	Country Specific Risk	IPEADATA
Consumer Confidence Index	ΔCCI	Consumer Confidence Index based on actual and future economic conditions, $2010 = 100$	General perception about Brazilian economy on both present and future	IInvestment Opportunities	FGV and Fecomercio/SP
Subsidized credit	ΔSUBS-GDP	Subsidized credit in the economy as a percentage of GDP	Reflects the magnitude of governamental subsidies to firms' investment	Subsidies	BCB
MSCI World	ΔMSCI-WD	MSCI World Index	Refers for the global equity market performance	Global Business Cycle	BCB
Fed Funds rate	ΔFED-FUNDS	US Federal Funds interest rate (%)	Proxy for the "flight-to-quality" movement of international investors	U.S. Short Term Interest Rates	IPEADATA
U.S. 10-Year Bond Yield	ΔUS10YBOND	U.S. 10-Year Bond Yield (%)	Proxy for the "flight-to-quality" movement of international investors	U.S. Long-Term Interest Rates	IPEADATA
2008 Financial Crisis	2008CRISIS	Dummy that equals 1 if 2008m6 $<= t <=$ 2008m11, and zero otherwise	Global financial crisis period	External shocks	Authors calculation
Brazilian Recession Periods	RECESSION	Dummy that equals 1 if there is a dated economic recession in Brazil, and zero otherwise	Local business cycle	Local business cycle	CODACE

Source: Authors' elaboration.

Analyzing the descriptive statistics of the variables (Table 9), we can see that the average monthly investment growth ($\Delta INVESTMENT$) during the full period (1996m3 - 2015m10) is 0.2%, while the median is 0.3%. The minimum value (-8.0%) observed in the variable occurred in 2013m12, when the Brazilian gross capital formation was already falling[31], despite the fact that the economic recession began only in 2014Q2. Meanwhile, the mean of net foreign portfolio investment-GDP ratio (EFPI-GDP) is 0.1%, ranging from -0.4% (2008m10) to 0.8% (2009m10). All other considered variables main statistics are reported in Table 9.

Table 9: Descriptive Statistics of the Variables

Note: This table reports the descriptive statistics of our variables for the 1996m3 - 2015m10 period. All variables shown here are stationary, according to the ADF tests. We use the following abbreviations: SD = Standard Deviation, Min = Minimum, Max = Maximum, and N = Number of observations.

Variable	**Mean**	**Median**	**Std Dev**	**Minimum**	**Maximum**	**N**
ΔINVESTMENT	0.002	0.003	0.015	-0.080	0.037	235
EFPI-GDP	0.001	0.000	0.001	-0.004	0.008	244
ΔRIR	-0.021	-0.130	2.474	-7.860	25.17	235
ΔSWAP360	-0.027	-0.040	1.665	-8.020	10.45	235
ΔCRED-GDP	0.088	0.110	0.467	-3.440	1.61	235
ΔSUBS-GDP	0.041	0.000	0.201	-0.200	2.000	235
ΔRER	0.002	0.000	0.042	-0.112	0.23	235
ΔIBOV	0.009	0.015	0.087	-0.503	0.215	235
ΔCCI	-0.001	0.002	0.051	-0.254	0.169	235
ΔEMBI-BR	-0.003	-0.017	0.116	-0.301	0.522	235
ΔCOMM	0.001	0.000	0.028	-0.169	0.082	235
ΔMSCI-WD	0.005	0.012	0.046	-0.210	0.107	235
ΔFED-FUNDS	-0.021	0.000	0.168	-0.960	0.280	238
ΔUS10YBOND	-0.020	-0.035	0.221	-1.110	0.650	238
2008CRISIS	0.013	0.000	0.112	0.000	1.000	236
RECESSION	0.267	0.000	0.443	0.000	1.000	236

Source: Authors' elaboration.

[31] Even though December is a typical month of low industry levels due to seasonality, our data is seasonally adjusted by the IBGE, and therefore this sharp decrease in aggregate investment is unlikely to be related to this issues. Nevertheless, this period coincides with the first quarter of decline in the Brazilian gross fixed capital formation, which has dropped impressively 25.0% from 2013Q3 (peak) to 2015Q4.

2.2.3 Stationarity tests

Stationarity is a key assumption for most of time series models. Although common practice in time series modeling has involved the application of Augmented Dickey Fuller (ADF) and Phillips-Perron (PP) tests to determine whether a series have a unit root, Elliot, Stock and Rothenberg (1996) show that a modified version of Dickey Fuller test using a generalized least squares (GLS) approach - named "DF-GLS" has substantially improved power when an unknown mean of trend is present. Using a 5% critical value cutoff, we find that most of our continuous variables are non-stationary in level, but they are stationary in first difference (i.e., they are I(1)). The exception is our measure of net foreign capital flows (EFPI-GDP), which is I(0). We also run some different specification tests and lag choosing[32], but the results hold for all series. We henceforth log-differentiate most of our variables - we use the "Δ" symbol to refer to them -, except for EFPI-GDP and our dummy variables (2008CRISIS and RECESSIONS), which are already stationary in levels.

2.2.4 Modeling the relationship between EFPI and aggregate investment

We consider the following vector autoregression with exogenous variables of order (p, q), denoted VARX$_{k,m}(p, q)$:

$$\mathbf{y}_t = \nu + A_1\mathbf{y}_{t-1} + \ldots + A_p\mathbf{y}_{t-p} + B_1\mathbf{x}_t + B_2\mathbf{x}_{t-1} + \ldots + B_q\mathbf{x}_{t-q} + \boldsymbol{\epsilon}_t \tag{5}$$

where $\{\mathbf{y}_t\}_{t=1}^T$ is an $N \times 1$ vector of endogenous variables and we allow for the presence of an $(M \times 1)$ vector of exogenous variables, $\{\mathbf{x}_t\}_{t=1}^T$. In equation (5), ν is an $N \times 1$ vector of intercepts, $A_j,\ j = 1, \ldots, p$ and $B_i,\ i = 1, \ldots, q$ are $N \times N$ and $N \times M$ matrices of slope coefficients, respectively, and $\epsilon_t \sim$ i.i.d. $(0, \Sigma_\epsilon)$ denotes a N-dimensional serially uncorrelated error term vector with mean zero and nonsingular covariance matrix Σ_ϵ[33], also referred to as a structural innovation or structural shock. A VAR, which is a special case of the VARX, can be represented by Equation (5) with $B_j = 0$ for $j = 1, \ldots, q$.

[32] We choose the optimum lag for the unit root tests based on the Akaike Information Criteria (AIC), being twelve the maximum allowed lag.

[33] For a moving-average representation of equation (5) to exist, the coefficient matrices A_j and B_j must be absolutely summable. This can be guaranteed, for example, by taking first differences of the endogenous variables.

Estimation of the parameters of the VAR in (5) requires that both Y_t and X_t are covariance stationary, i.e., have their first two moments finite and time-invariant. It will be convenient to represent the model by its Vector Moving Average (VMA) form[34]

$$\mathbf{y}_t = \mu + \sum_{i=0}^{\infty} \Lambda_i \epsilon_{t-i} + \sum_{i=0}^{\infty} \Psi_i \mathbf{x}_{t-i} \tag{6}$$

where the endogenous variables are expressed as a function of a constant N-vector (μ), and the current and past values of the structural shocks (ϵ) and the exogenous variables. The Ψ_i matrices are the dynamic multiplier functions (or transfer functions), and the sequence of moving average coefficients Λ_i are the simple impulse-response functions (IRFs) at horizon i.

In our specific setting, we consider the vector of endogenous variables $\mathbf{y}_t = [\Delta\mathsf{INV}_t, \Delta\mathsf{RIR}_t, \Delta\mathsf{RER}_t, \Delta\mathsf{IBOV}_t, \mathsf{EFPI}_t]$, where INV_t is the aggregate investment, RIR_t is the real interest rate, RER_t is the real exchange rate, IBOV_t is the Brazilian stock market index called IBOVESPA, and EFPI_t is the equity foreign portfolio investment normalized by GDP. Because the residual covariance matrix Σ_ϵ is generally not diagonal (i.e., the components of u_t are contemporaneously correlated[35]), we propose a recursive identification such that:

$$\begin{pmatrix} \epsilon_{t,\Delta\mathsf{INV}} \\ \epsilon_{t,\Delta\mathsf{CRED\text{-}GDP}} \\ \epsilon_{t,\Delta\mathsf{RIR}} \\ \epsilon_{t,\Delta\mathsf{RER}} \\ \epsilon_{t,\Delta\mathsf{IBOV}} \\ \epsilon_{t,\mathsf{EFPI}} \end{pmatrix} = \begin{bmatrix} \alpha_{11} & 0 & 0 & 0 & 0 & 0 \\ \alpha_{21} & \alpha_{22} & 0 & 0 & 0 & 0 \\ \alpha_{31} & \alpha_{32} & \alpha_{33} & 0 & 0 & 0 \\ \alpha_{41} & \alpha_{42} & \alpha_{43} & \alpha_{44} & 0 & 0 \\ \alpha_{51} & \alpha_{52} & \alpha_{53} & \alpha_{54} & \alpha_{55} & 0 \\ \alpha_{61} & \alpha_{62} & \alpha_{63} & \alpha_{64} & \alpha_{65} & \alpha_{66} \end{bmatrix} \begin{pmatrix} u_{t,1} \\ u_{t,2} \\ u_{t,3} \\ u_{t,4} \\ u_{t,5} \\ u_{t,6} \end{pmatrix} \tag{7}$$

Where each line can be viewed as an equation (multiplying through each term on the right-hand side), and each reduced-form shock is a weighted average of selected structural shocks. The triangular matrix described in 11 is the Cholesky Decomposition of the variance-covariance matrix

[34] The existence of this representation is ensured if the VAR process is stable, i.e., if y_t consists of stationary (I(0)) variables. Intuitively, the system's stability ensures that the effect of an impulse in a variable in the system is transitory (see, for example, LÜTKEPOHL, 2011 and BAUM, 2013).

[35] Because of space constraints, we do not present the covariance matrix of the estimated residuals. However, they seem to be – as expected – correlated across equations (e.g., the correlation between the residuals of the real interest rate and credit-GDP ratio equations is .1144349).

of residuals, and it ensures that the error terms on each equation in the system are orthogonal.[36] This "triangularization" of the VAR was first suggested by Sims (1980), and it became widely used by macroeconometricians over the past decades. Our identification strategy resides in the establishment of an order among system variables that is economically plausible, which meets the most recent recommendations in the macro-finance literature (see, e.g., STOCK; WATSON, 2001; LÜTKEPOHL, 2010; ROMER, 2012).

Our strategy allows tracing the effects of a shock in variable "j" in time "t" (u_{jt}) in all variables in the VARX system, keeping all else equal. The short-run restrictions we impose in the coefficients in order to obtain identification – also known as zero restrictions – is also corroborated by the frequency of our data: if one had annual data, a contemporaneous restrictions would likely be more debatable that if it were on a quarterly or monthly basis (RONAYNE, 2011). In this kind of VAR study, the classical identification problem (correlation vs causation) cannot be solved by a purely statistical tool - rather, economic theory or institutional knowledge must be used to solve the identification issue (see, e.g., STOCK; WATSON, 2001; ROMER, 2012). Even though we try to rely on economic models, our identification scheme does not build on the "natural experiment" literature, which is still a big open challenge for macro-economists (ROMER, 2012).

2.2.4.1 Exogeneity and marginal significance of variables

To complement the endogenous variables defined in the VAR system, we add variables that are plausibly exogenous to our model, i.e., variables that represent external conditions or the global business cycle, and probably are not (at least largely) affected by the Brazilian economy's behavior. Our motivation is based on the perception that external conditions, such as global liquidity, uncertainty and the global business cycle may affect capital inflows to emerging markets, which is supported by the empirical evidence that point out global factors as key determinants to net capital inflows to developing economies (e.g., CALVO; LEIDERMAN; REINHART, 1996).

Specifically, we consider the MSCI World Index as a proxy for the world equity market performance. We also add the Fed Funds rate as a proxy for the "flight-to-quality" movement of international investors, and we expect the opportunity cost of investing in emerging markets to

[36] If the covariance matrix of structural shocks Σ_ϵ is diagonal, the structural shocks are assumed to be uncorrelated. In our VARX system, it means that, e.g., a surprise effect in the equity foreign portfolio investment can affect aggregate domestic investment in the same month, but not the other way around.

be negative related with capital inflows to Brazil. Because of the impact on terms-of-trade, we also consider the IMF Primary Commodity Prices index (ex-fuel, which is fairly approximated to Brazilian net exporting[37]) to capture the terms-of-trade shocks to the local economy. Finally, we consider the monthly chronology of recessions in the Brazilian economy dated by the Business Cycles Dating Committee (CODACE), that accounts for different economic regimes in the business cycle. Finally, we test the jointly significance of the coefficients of these variables in our k=6 equations in the unrestricted VARX system (Table 10). As parsimonious is a key feature here, we only keep in our models variables that are marginally significant in our models - we consider a 0.1 p-value trashold to determine whether a given variable should be included or not in the estimated model.

Table 10: Joint test of significance of the exogenous variables coefficients

Note: This table reports the test statistics of the marginal significance of each exogenous variable in the estimated VARX system. Prob > Chi2 reveals the p-value associated with the null hyphotesis that all lags of the exogenous variables are equal to zero in all system's equations.

Variable	**Chi2**	**Prob > Chi2**
L2.ΔCOMM	10.89	0.0920
ΔMSCI-WD	209.69	0.0000
ΔFED-FUNDS	1.81	0.9367
ΔUS10YBOND	2.68	0.8483
2008CRISIS	11.76	0.0676
RECESSION_DUM	28.87	0.0001

Source: Authors' elaboration.

Complementing our analysis of variables that are potentially being determined outside the model we analyze, we also perform some bivariate Granger Causality Tests in order to check if the lags of our dependent variable (investment) do not help to explain current values of the exogenous variables. Indeed, most of these variables representing external conditions helps to explain the endogenous variable meanwhile they are not affected by them (there is a unidirectional Granger-Causality from exogenous to endogenous variables). For those reasons, we keep in the model only the plausibly exogenous variables that are marginally significant in our models and are not

[37] Brazil's main exports include soybeans, coffee, tobacco, cocoa, beef, poultry, orange juice, raw cane and refined sugar, iron ore and concentrates, oil seed, and mineral fuels.

Granger-caused by the endogenous variables in our model.[38]

Table 11: Table: LM test for residual autocorrelation in the model

Note: This table reports the statistics of the LM test for residual autocorrelation after the VAR. In the first model (A), maximum lag is set to be equal to the optimal lag-length suggested by the Akaike Information Criteria (AIC). Since there is evidence of autocorrelation in at least one lag (p<0.10), we test in the second model (B) the same specification with one extra lag for all endogenous variables (INVESTMENT, CRED-GDP, RIR, RER, IBOV and EFPI). Prob>Chi2 reveals the probability of rejecting the null hypothesis that there is no residual autocorrelation in that specific lag. the test statistics of the marginal significance of each exogenous variable in the estimated VAR system. Prob > Chi2 reveals the p-value associated with the null hypothesis that all lags of the exogenous variables are jointly equal to zero.

A) Maximum Lag = 2			
Lag	Chi2	DF	Prob > Chi2
1	48.7194	36	0.07659
2	47.6566	36	0.09256
3	34.8732	36	0.52206
B) Maximum Lag = 3			
Lag	Chi2	DF	Prob > Chi2
1	37.7648	36	0.38858
2	40.1045	36	0.29308
3	33.6176	36	0.58242

Source: Authors' elaboration.

2.2.4.2 Lag selection and residual autocorrelation

We use information criteria to define the lag-length of our VAR models. Simulations show that for monthly VAR models, the AIC tends to produce the most accurate impulse response estimates independent of the sample size (IVANOV; KILIAN, 2005). For that reason, we choose the AIC criterion instead of the HQIC, SBIC or other available lag-length selection criteria. Because our data is monthly, we set the maximum lag-length to twelve and check the AIC results for all different lag-lengths, trying to find the minimum AIC value. Both the AIC and Final Prediction Error (FPE) suggest that the optimum lag is equal to two, while HQIC and SBIC recommended lag equal to one.

Following the optimum lag choice (p*=2 as sugested by AIC), we proceed to a model

[38] The exception from this analysis is the 2008CRISIS dummy. Although this variable is marginally significant in the model (see Table 10), its periods coincides with the 2008-2009 economic recession in Brazil dated by CODACE, which is already included in the model. Moreover, since we split our sample in before and after the 2008 financial crisis, the 2008CRISIS dummy would be naturally out of the model in all tests using this sub-sample.

checking of its adequacy (LÜTKEPOHL, 2011). Because residual autocorrelation may lead to inneficient cofficient estimations using OLS, we perform the Breusch-Godfrey Lagrange Multiplier (LM) test to check if the residuals are white noise, i.e., if they have zero mean and constant variance. According to Lütkepohl (2011), this test is recommended for low order autocorrelation, instead of the Portmanteau test, which is recommended for high-order correlation. Specifically, the LM test for residual autocorrelation is a test for zero coefficient matrices for the residuals,

$$u_t = B_1 u_{t-1} + ... + B_h u_{t-h} + e_t \tag{8}$$

under the null hypothesis that all coefficients of the past errors are jointly equal to zero. Specifically, we test

$$H_0 : B_1 = ... = B_h = 0 \text{ versus } H_1 : B_i \neq 0 \text{ for at least one } i \in 1, ..., h.$$

As one could note from Table 11, the LM test suggest that there is a significant residual autocorrelation at the optimal lag-length suggested by AIC (maximum lag=2), so we add one more lag and test for residual autocorrelation again. Table 11 reveals that we end up getting white noise residuals in the model with lag*=3. Using this rule of thumb, the model with lag = 3 is sufficient to eliminate the autocorrelation of the residuals, at the same time that the model does not become superparametrized. Therefore, we use three lags of each endogenous variables to estimate our VARX systems.

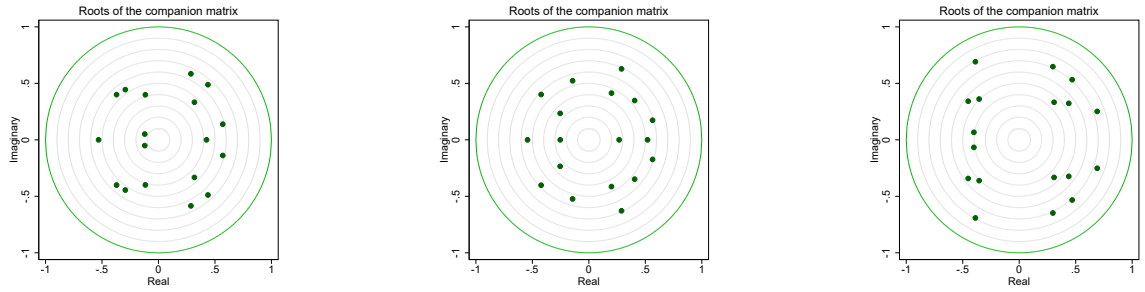
2.2.5 Model Stability

One of the key assumptions for generating IRFs is that the VAR must be stable (BAUM, 2013). The stability of the system is complementary to the assumption of covariance stationary for the variables, i.e., that variables should have a constant mean and a constant variance. If the model is stable, it has an infinite-order vector moving-average representation, and thus impulse-response functions and forecast-error variace decompositions have known interpretations (STATA, 2015). Figure 7 resumes the stability analysis for the full period (a), and also for our splited sample in before (b) and after the 2008 financial crisis (c).

As pointed out by Lütkepohl (2011), if the modulus of each eigenvalue of the matrix A is

Figure 7: Stability of the VARX System - Full period and before/after the 2008 financial crisis

(a) Full period (1996m3 - 2015m10) (b) Before crisis (1996m3 - 2008m8) (c) After crisis (2008m12 - 2015m10)



Source: Authors' elaboration.

strictly less than one, then the estimated VAR is stable. In our case, we can infer from Figure 7 that the eigenvalues are inside the unit circle, indicating that the system is stable. In other words, all variables in the system are jointly stationary and our model is "non-explosive".

2.3 Results

In this subsection, we exploit the main results of our analysis. We start by looking at the VAR coefficients, further analyzing Granger Causality, Impulse-Response functions, and FEVDs.

2.3.1 VAR coefficients

We begin our analysis by looking at a model estimated for the full period covered in our study (1996-2015). The results of the estimation of the XVAR model are shown in Table 12. According to the model specification, gross capital formation ($\Delta INVESTMENT$), credit as a share of GDP ($\Delta CRED - GDP$), real interest rates (ΔRIR), real exchange rates (ΔRER), stock market valuation ($\Delta IBOV$) and net equity foreign portfolio investments (EFPI-GDP) are simultaneously and endogenously determined in the VAR system. Additionally, a set of exogenous variables also affect the endogenous variables: commodity prices ($\Delta L2.COMM$), current global stock market performance ($\Delta MSCI - WD$), and a dummy capturing periods of recessions in the economy (RECESSION). Our analysis will focus on columns 2 ($\Delta INVESTMENT$) and 4 (EFPI-GDP), but we will comment on the effects of foreign capitals on other variables whenever such effects are relevant.

Table 12: Vector Autoregressive Model - Full period (1996-2015)

Note: This table presents the results of the VARX System of simultaneous equations for the period 1996m3-2015m10. Dependent variables are listed in the columns ($\Delta INVESTMENT$, $\Delta CRED-GDP$, ΔRIR, ΔRER, $\Delta IBOV$, and EFPI-GDP) and represent each equation of the dynamic system. We select the optimum lag length according to AIC plus testing for autocorrelation in residuals using a Lagrange Multiplier (LM) test. If residuals are autocorrelated, we include one extra lag to the model until we do not reject the null hypothesis that there is no residual autocorrelation in the model (our optimum lag length following this criteria is $lag = 3$). The estimated coefficients and the t statistics (in parentheses) are reported for each independent variable. ***, **, and * implies significance at the 99%, 95% and 90% levels, respectively.

Variable	$\Delta INVESTMENT$	$\Delta CREDIT-GDP$	RIR	ΔRER	$\Delta IBOV$	EFPI-GDP
L.$\Delta INVESTMENT$	0.787***	2.098	3.618	−0.405*	−0.177	−0.002
L2.$\Delta INVESTMENT$	−0.341***	6.110*	18.065	0.443	−0.396	0.005
L3.$\Delta INVESTMENT$	−0.030	−2.608	−7.948	−0.216	0.062	0.006
L.$\Delta CRED-GDP$	−0.001	0.111***	−0.548	−0.005	0.008	0.000
L2.$\Delta CRED-GDP$	−0.002	−0.005	0.298	0.015***	0.002	−0.000
L3.$\Delta CRED-GDP$	0.001	0.111*	0.273	0.001	0.001	0.000
L.ΔRIR	−0.000	-0.072***	0.109*	0.001	0.003*	−0.000
L2.ΔRIR	0.000	0.017	-0.167**	−0.001	-0.003*	0.000
L3.ΔRIR	0.000	0.013	−0.033	−0.000	−0.003	0.000
L.ΔRER	−0.008	−0.618	−7.067*	0.317***	0.079	−0.001
L2.ΔRER	−0.015	−0.565	−0.925	−0.104	0.088	0.003
L3.ΔRER	−0.004	−0.524	−2.553	−0.009	−0.150	−0.002
L.$\Delta IBOV$	0.005	−0.555	-9.934***	0.010	−0.032	0.004***
L2.$\Delta IBOV$	0.000	−0.605	2.128	−0.046	0.088	0.001
L3.$\Delta IBOV$	0.008	0.085	-3.525*	-0.088***	−0.053	−0.000
L.EFPI-GDP	1.126*	−3.559	−127.815	−2.549	0.280	0.021
L2.EFPI-GDP	−0.256	−3.996	−124.021	-4.270*	−2.686	0.106
L3.EFPI-GDP	0.648	−40.926	125.756	1.905	−6.571*	0.088
$\Delta L2.COMM$	0.061**	0.558	−6.092	-0.223**	0.011	0.005*
$\Delta MSCI-WD$	0.005	-2.049***	3.395	0.069	1.336***	0.009***
RECESSION-Dummy	-0.007***	−0.069	0.328	0.000	−0.015	−0.000
CONSTANT	0.002**	0.130***	0.024	0.005	0.011*	0.000***
RMSE	0.009	0.422	2.292	0.0347	0.063	0.001
$\overline{R}^2$	0.670	0.258	0.215	0.371	0.530	0.302
χ^2	471.78	80.63	63.58	136.96	262.23	100.53
$P > \chi^2$	0.000	0.000	0.000	0.000	0.000	0.000
Observations	232	232	232	232	232	232

Source: Authors' elaboration.

In the second column of Table 12, the coefficients for $\Delta INVESTMENT$ (our variable of main interest) are reported. In the full period analysed, the model reveals a strong dependence between current investment growth and past movements of investment growth (+), in line with theories of irreversible investment. We find that real interest rate, a proxy for the real rental price of capital, does not significantly affect changes in the investment levels. This evidence is consistent with other studies evaluating aggregate investment in Brazil (e.g., LUPORINI; ALVES, 2010), and suggest that investment responses to past movements in real interest rates are close to zero. Our model also rejects any effects of the supply of credit and exchange rate variations on investment,

but we find that investment is affected by past movements of EFPI-GDP (+) and L2.$\Delta COMM$ (+), and also contemporaneously by RECESSION (-).

As shown in our VAR model, foreign portfolio capital flows (EFPI-GDP) exerts a positive and statistically significant effect on investment growth. This finding is consistent with theories of financial integration, as foreign capital flows improve risk sharing between domestic and foreign investors, driving the cost of equity capital downwards and hence increasing investment (STULZ, 2005). Such result is also in line with prior empirical work, corroborating the argument that foreign capitals may cause investment booms and contribute to economic growth (HENRY, 2000; BEKAERT; HARVEY; LUNDBLAD, 2005). Other variables which entered the model exogenously also help explaining the investment behavior. Investment grows when global commodity prices are high, which is reasonable given the dependence of the Brazilian economy on commodity exports (especially agricultural commodities and basic materials), thus it makes sense that firms increase investments when commodity prices are soaring, for their expected revenues increase.[39] Moreover, investment slows when the economy is facing a recessive period, which follows naturally from standard macroeconomic models.

In column 3, the estimated coefficients for the real exchange rate (ΔRER) equation are shown. Foreign capitals affect the real exchange rate between the Brazilian Real and the U.S dollar (-). We interpret this finding as an outcome from the interplay between supply and demand for foreign currency in the domestic exchange market. As foreign capitals cause a greater influx of foreign currency, for foreign investors must buy Brazilian currency to purchase local equities, the relative value of the Brazilian currency appreciates. Such empirical evidence of a negative effect of foreign capitals on the exchange rate might be seen as a justification backing governmental intervention by means of capital controls, which took place especially between years 2009 and 2011 (CHAMON; GARCIA, 2016; JINJARAK; NOY; ZHENG, 2013).

In column 4, the estimation results of the stock market valuation equation ($\Delta IBOV$) are shown. As stock market valuation enters the equation in log differences, coefficients read as partial effects on stocks returns. Foreign equity capital affects stock market returns (-), and the effect occurs with a three periods lag. Such a negative effect of foreign equity capital on expected returns

[39] In a recent paper, Shousha(2015) finds that commodity prices shocks are an important source of business cycle fluctuations for emerging markets, especially those classified as net commodity exporters, such as Brazil.

is fully in line with the argument that financial integration brought about by foreign equity capital investments improve risk sharing between foreign and domestic investors, reducing the cost of equity capital (BUCKBERG, 1995; HENRY, 2000; BEKAERT; HARVEY; LUNDBLAD, 2005; DE JONG; DE ROON, 2005; CARRIERI; ERRUNZA; HOGAN, 2007).

The last column of Table 12 shows the results for the equation in which we model the determinants of foreign capital flows (EFPI-GDP). The main determinant of foreign capital flows is $\Delta IBOV$ (lagged stock market valuation) (+). This empirical finding corroborates theories of foreign investor behavior, namely the so-called positive feedback trading hypothesis, which argues that foreign investors adopt momentum strategies, investing when expected capital gains are high due to increases in equity valuations (FROOT; O'CONNEL; SEASHOLES, 2001; KAMINSKY; LYONS; SCHMUKLER, 2004). The variable $\Delta MSCI - WD$, which captures global stock market performance, also affects foreign capital flows (+), hence the Brazilian economy receives more foreign capital when the outlook in global markets is positive. Finally, foreign capitals are influenced by global commodity prices (+), which possibly reflects higher equity valuations of Brazilian stocks when commodity prices soar.

In Table 13 we show the results for the VAR models specified on two different sampling periods. The first period relates to the pre-crisis period, and runs from 1996 up to the eve of the crisis, August, 2008, the month in which the Brazilian equity market was first hit by the global turmoil. The second period goes from December, 2008, the month that marks the initial recovery in the Brazilian equity market, up to the end of our sampling period, in October, 2015. Panel A refers to the pre-crisis period, whereas Panel B to the post-crisis period.

This analysis shows that the positive effect of foreign capitals on investment growth refers to the pre-crisis period. Comparing the full period with the pre-crisis sub-period, we see that during the pre-crisis period the coefficient is slightly larger (1.29 vs 1.12), and the statistical significance is stronger ($p<0.05$ vs $p<0.10$). In fact, in the post-crisis period, the effect of of foreign capitals on investment growth is not statistically different from zero. Thus, it seems that the 2008 financial crisis has somehow interrupted or severely affected the ability of foreign capitals to fund the domestic capital stock. We discuss possible reasons behind this shift later on in the paper.

In general, the empirical evidence backed by the coefficients fitted by our VAR analysis

provides support for an active role of foreign capitals in financing investment growth. However, the model coefficients and the R^2 statistics are a first evidence of the effects of each variable in the system only. We follow Stock and Watson (2001) and revisit these relationships employing more robust analytic techniques. Because of the complicated dynamics in the VAR, Granger-causality tests, impulse response functions and forecast error variance decompositions are more informative than the estimated VAR regression coefficients. For that reason, we limit our analysis of the model coefficients in order to focus on the more informative analysis of the results. We conduct such analyses in the next two sections.

Table 13: VAR Models - Two sub-periods (pre and post 2008 financial crisis)

Note: This table presents the results of the VAR System of simultaneous equations for our splited sample (1996m3 - 2008m8 and 2008m12 - 2015m10). Dependent variables are listed in the columns ($\Delta INVESTMENT$, $\Delta CRED - GDP$, ΔRIR, ΔRER, $\Delta IBOV$, and EFPI-GDP) and represents each equation of the dynamic system. We select the optimum lag length according to AIC plus testing for autocorrelation in residuals using a Lagrange Multiplier (LM) test. If residuals are autocorrelated, we include one extra lag to the model until we do not reject the null hyphotesis that there is no residual autocorrelation in the model (our optimum lag length following this criteria is $lag = 3$). The estimated coefficients and the t statistics (in parentheses) are reported for each independent variable. ***, **, and * implies significance at the 99%, 95% and 90% levels, respectively.

Variable	PANEL A: Before the financial crisis (1996m3 - 2008m08)						PANEL B: After the financial crisis (2008m12 - 2015m10)					
	$\Delta INVESTMENT$	$\Delta CRED - GDP$	ΔRIR	ΔRER	$\Delta IBOV$	EFPI-GDP	$\Delta INVESTMENT$	$\Delta CRED - GDP$	ΔRIR	ΔRER	$\Delta IBOV$	EFPI-GDP
L.$\Delta INVESTMENT$	0.768^{***}	6.864	10.438	-0.221	0.053	-0.007	0.724^{***}	-2.294	1.008	-0.616^{**}	0.293	0.015
L2.$\Delta INVESTMENT$	-0.211^{**}	-1.964	33.981	0.501	-0.796	0.019	-0.336^{**}	9.159^{***}	1.599	0.413	-0.571	-0.007
L3.$\Delta INVESTMENT$	0.018	1.928	-13.321	-0.395	0.178	0.000	-0.241^{**}	0.481	4.303	-0.443	-0.059	0.013
L.$\Delta CRED - GDP$	-0.002^{**}	0.093	-0.593	-0.005	0.012	0.000	0.010^{**}	-0.074	-0.217	-0.012	-0.003	0.001^{**}
L2.$\Delta CRED - GDP$	-0.001	-0.005	0.532	0.010	0.003	-0.000	0.004	-0.130	-0.049	0.027^{**}	0.005	0.001^{*}
L3.$\Delta CRED - GDP$	0.002	0.012	0.414	0.003	0.002	0.000	0.002	0.441^{***}	-0.106	-0.020^{**}	0.002	0.001
L.ΔRIR	-0.000^{**}	-0.071^{***}	0.080	0.001	0.003	-0.000	0.006^{**}	-0.005	0.513^{***}	0.000	0.005	-0.000
L2.ΔRIR	0.000	0.018	-0.173^{**}	-0.001	-0.004^{*}	-0.000	-0.003	-0.036	0.058	0.009	-0.008	0.000
L3.ΔRIR	0.000	0.017	-0.030	-0.001	-0.003^{*}	0.000	-0.002	-0.026	-0.129	-0.016^{**}	0.001	0.001
L.ΔRER	-0.011	-0.913	-9.541	0.441^{***}	0.278^{*}	-0.001	-0.074^{*}	-1.146	-0.512	0.083	-0.162	0.000
L2.ΔRER	-0.005	-0.356	-1.036	-0.154^{*}	0.015	0.003	-0.064^{*}	-1.142	-0.642	-0.074	0.241	0.002
L3.ΔRER	-0.016	-0.872	-5.510	0.054	-0.031	-0.003	-0.021	0.609	1.577	-0.099	-0.111	-0.000
L.$\Delta IBOV$	-0.008	-0.562	-11.986^{***}	0.014	-0.068	0.003^{***}	0.055^{***}	-0.348	-0.793	0.037	0.029	0.007^{***}
L2.$\Delta IBOV$	-0.003	-0.532	2.759	-0.032	0.081	0.001	0.017	0.720	-0.845	-0.148^{***}	0.003	0.001
L3.$\Delta IBOV$	0.015^{**}	-0.089	-2.740	-0.121^{***}	-0.095	-0.000	-0.017	0.874^{*}	-0.824	-0.027	-0.064	0.001
L.EFPI-GDP	1.292^{**}	-11.534	-306.246	-1.765	0.121	-0.044	-0.072	-31.996	-22.319	-0.932	4.601	-0.059
L2.EFPI-GDP	-0.689	-15.539	-332.164	-1.380	-3.045	0.099	-0.348	-27.027	11.919	-4.610^{**}	2.540	0.029
L3.EFPI-GDP	0.997	-5.054	98.566	2.100	-4.687	0.059	-0.815	-73.349^{***}	25.664	2.536	-2.932	0.055
L2.$\Delta COMM$	-0.011	2.239	-15.398	-0.012	0.204	0.001	0.128^{***}	-0.362	-1.241	-0.363^{***}	-0.244	0.010^{**}
$\Delta MSCI - WD$	0.004	-2.223^{**}	2.973	0.100	1.711^{***}	0.007^{***}	-0.005	-1.164^{*}	0.005	0.064	0.898^{***}	0.005
RECESSION	-0.006^{***}	-0.132	0.548	0.000	-0.028^{**}	-0.000	-0.009^{***}	0.054	0.035	-0.001	0.002	0.000
CONSTANT	0.002^{**}	0.090	0.230	0.003	0.019^{***}	0.000^{***}	0.002	0.212^{***}	0.028	0.006	-0.011	0.000
$\overline{R}^2$	0.657	0.279	0.273	0.329	0.580	0.245	0.774	0.469	0.440	0.691	0.560	0.418
RMSE	0.007	0.491	2.856	0.038	0.069	0.001	0.010	0.250	0.356	0.025	0.044	0.001
Observations	146	146	146	146	146	146	83	83	83	83	83	83

Source: Authors' elaboration.

2.3.2 A closer look at causality

An usual approach to examine the VAR results is to proceed with Granger-Causality tests after the model is fitted (STOCK; WATSON, 2001). Basically, a variable x is said to Granger-cause a variable y if, given the past values of y, including past values of x is useful to predict y. The way we test if each endogenous variable in the system Granger-Cause others is to regress y_i, $i = 1, ..., 6$ on its own lags (lag* = 3) and on lagged values of other variables. The null hypothesis of the test is that all estimated coefficients of the lagged values of x are jointly zero. Failure to reject the null hypothesis is equivalent to failing to reject the hypothesis that x does not Granger-cause y. Granger causality test for all endogenous variables in the VAR system are shown in Table 14.

Again, we focus our analysis on the determinants of aggregate investment growth. In the first column of Table 14, we show Granger causality tests for the full period covered in our study. The model does not do so well in explaining the determinants of the growth in capital formation, at least not for the full period taken altogether, as we do not find any statistical significant causal relationships between investment growth and the other endogenous variables in the model.

However, when looking at the pre and post-crisis sub-periods, we see causalities emerging. In the pre-crisis period, real interest rates (-), foreign capital flows (+) and credit supply (-) Granger-caused investment growth, as the hypothesis of non-causality is rejected for the three variables, at 95% confidence level for EFPI-GDP and at 90% level for ΔRIR and $\Delta CRED - GDP$. Also, the causality test for all variables jointly rejects the null hypothesis of non-causation at 99% confidence level.

In the post-crisis period, we see that stock market returns take on more importance in explaining investment growth (+), as $\Delta IBOV$ granger causes investment, and we also find a statistically significant causality of exchange rate variations on investment (-). Again, the joint test of causality for all variables taken together rejects the null hypothesis of non-causation at 99% confidence level. In line with the findings from the VAR coefficient analysis, we do not find any causality running from foreign capitals to investment in the post-crisis period, what reinforces the argument that the role of foreign equity capital in financing investment has changed somehow, interrupting a very beneficial economic relationship between the two variables.

In line with the previous results found in the analysis of the VAR coefficients, we see that

Table 14: Granger Causality Tests

Note: This table presents the results of the Granger-causality tests conducted after the estimation of the VAR System of simultaneous equations. We test if each endogenous variable in the system Granger-Cause others by regressing y_i, $i = 1, ..., 6$ on its own lags (lag* = 3) and on lagged values of other variables. The null hypothesis of the test is that all estimated coefficients of the lagged values of x are jointly zero. Failure to reject the null hypothesis is equivalent to failing to reject the hypothesis that x does not Granger-cause y. ***, **, and * implies significance at the 99%, 95% and 90% levels, respectively. All variables in tests are stationary (we suppress the "Δ" symbol because of space constraints, but it is applied to INVESTMENT, CRED-GDP, RIR, RER, and IBOV).

Full period			Before 2008 crisis			After 2008 crisis		
Equation	Excluded	Prob	Equation	Excluded	Prob	Equation	Excluded	Prob
INVESTMENT	CRED-GDP	0.418	INVESTMENT	CRED-GDP	0.073	INVESTMENT	CRED-GDP	0.113
INVESTMENT	RIR	0.375	INVESTMENT	RIR	0.055	INVESTMENT	RIR	0.183
INVESTMENT	RER	0.628	INVESTMENT	RER	0.507	INVESTMENT	RER	0.054
INVESTMENT	IBOV	0.719	INVESTMENT	IBOV	0.148	INVESTMENT	IBOV	0.019
INVESTMENT	EFPI-GDP	0.112	INVESTMENT	EFPI-GDP	0.036	INVESTMENT	EFPI-GDP	0.811
INVESTMENT	ALL	0.174	INVESTMENT	ALL	0.002	INVESTMENT	ALL	0.007
CRED-GDP	INVESTMENT	0.025	CRED-GDP	INVESTMENT	0.427	CRED-GDP	INVESTMENT	0.000
CRED-GDP	RIR	0.000	CRED-GDP	RIR	0.000	CRED-GDP	RIR	0.828
CRED-GDP	RER	0.457	CRED-GDP	RER	0.513	CRED-GDP	RER	0.226
CRED-GDP	IBOV	0.181	CRED-GDP	IBOV	0.486	CRED-GDP	IBOV	0.139
CRED-GDP	EFPI-GDP	0.461	CRED-GDP	EFPI-GDP	0.980	CRED-GDP	EFPI-GDP	0.001
CRED-GDP	ALL	0.000	CRED-GDP	ALL	0.000	CRED-GDP	ALL	0.000
RIR	INVESTMENT	0.534	RIR	INVESTMENT	0.508	RIR	INVESTMENT	0.181
RIR	CRED-GDP	0.319	RIR	CRED-GDP	0.369	RIR	CRED-GDP	0.485
RIR	RER	0.279	RIR	RER	0.248	RIR	RER	0.495
RIR	IBOV	0.000	RIR	IBOV	0.000	RIR	IBOV	0.298
RIR	EFPI-GDP	0.551	RIR	EFPI-GDP	0.380	RIR	EFPI-GDP	0.762
RIR	ALL	0.000	RIR	ALL	0.000	RIR	ALL	0.305
RER	INVESTMENT	0.393	RER	INVESTMENT	0.742	RER	INVESTMENT	0.098
RER	CRED-GDP	0.033	RER	CRED-GDP	0.393	RER	CRED-GDP	0.004
RER	RIR	0.658	RER	RIR	0.394	RER	RIR	0.227
RER	IBOV	0.020	RER	IBOV	0.016	RER	IBOV	0.015
RER	EFPI-GDP	0.123	RER	EFPI-GDP	0.855	RER	EFPI-GDP	0.102
RER	ALL	0.000	RER	ALL	0.010	RER	ALL	0.000
IBOV	INVESTMENT	0.485	IBOV	INVESTMENT	0.702	IBOV	INVESTMENT	0.469
IBOV	CRED-GDP	0.869	IBOV	CRED-GDP	0.734	IBOV	CRED-GDP	0.990
IBOV	RIR	0.014	IBOV	RIR	0.019	IBOV	RIR	0.954
IBOV	RER	0.375	IBOV	RER	0.203	IBOV	RER	0.305
IBOV	EFPI-GDP	0.342	IBOV	EFPI-GDP	0.866	IBOV	EFPI-GDP	0.542
IBOV	ALL	0.109	IBOV	ALL	0.249	IBOV	ALL	0.804
EFPI-GDP	INVESTMENT	0.384	EFPI-GDP	INVESTMENT	0.247	EFPI-GDP	INVESTMENT	0.400
EFPI-GDP	CRED-GDP	0.834	EFPI-GDP	CRED-GDP	0.732	EFPI-GDP	CRED-GDP	0.035
EFPI-GDP	RIR	0.875	EFPI-GDP	RIR	0.840	EFPI-GDP	RIR	0.114
EFPI-GDP	RER	0.424	EFPI-GDP	RER	0.419	EFPI-GDP	RER	0.973
EFPI-GDP	IBOV	0.000	EFPI-GDP	IBOV	0.001	EFPI-GDP	IBOV	0.015
EFPI-GDP	ALL	0.038	EFPI-GDP	ALL	0.058	EFPI-GDP	ALL	0.027

Source: Authors' elaboration.

movements in stock market returns ($\Delta IBOV$) Granger-cause foreign capital flows (+), reinforcing the evidence on active feedback trading strategies put in place by foreign investors in the Brazilian

equity market. Interestingly, when revisiting our analysis of causalities our findings show once again that foreign capitals cause exchange rate depreciation, but after the crisis only, as we reject the hypothesis that foreign capitals do not Granger-cause exchange rate variations with 90% confidence interval. This additional piece of evidence corroborates the findings from the analysis of VAR coefficients. Moreover, we do not find any evidence suggesting that investment growth Granger-causes foreign capital flows, hence we can safely state that the causality runs from foreign capitals to investment growth and not the converse.

2.3.3 Impulse-Response Functions (IRFs)

To understand the dynamic properties of domestic investment, we follow the guidelines of Stock and Watson (2001) and compute impulse response functions (IRFs) for the foreign equity capital. A IRF traces the impact of a one-time, unit shock to one variable on the current and future values of the endogenous variables. Since the innovations are correlated (as we shall show), they are orthogonalized.[40] When computing the IRF, we need to choose a specific ordering of the endogenous variables since different orderings may result in different responses.[41]

Impulse-response functions for the recursive VARX are plotted in Figure 8. As our main variable of interest is the aggregate investment, each graph in the figure reflects the cummulative effect of an unexpected one unit increase in one endogenous variable(k=1, 2, ..., 6) on $\Delta INVESTMENT$, from one to fifteen months ahead. Following Stock and Watson (2001), we report the point estimates of the COIRFs accompanied by one asymptotic standard error band for each impulse response.

During the full period, Figure 8 indicates that a 1% shock to domestic investment growth lead to an approximately 1.5% cumulative impact on investment growth, with the response decaying rapidly from month one to month two and more gradually after that. Shocks to the foreign equity capital and to stock market returns also increase investment, but these effects are much weaker (around 0.4%) when compared to shocks to lagged investment. The response of investment

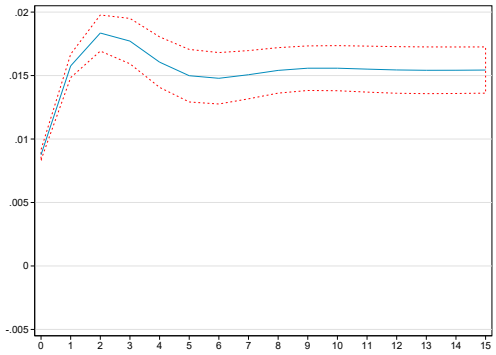
[40] Specifically, the inverse of the Cholesky decomposition factor of the residual variance-covariance matrix is used to orthogonalize the impulses. This procedure is important to a "ceteris paribus" analysis: if the residuals are correlated across equations, a shock to one variable will be confounded by the reaction in the other error terms, and therefore we can not have a causal interpretation.

[41] However, the VAR coefficient estimates and the Granger causality results are unaffected by the ordering of variables.

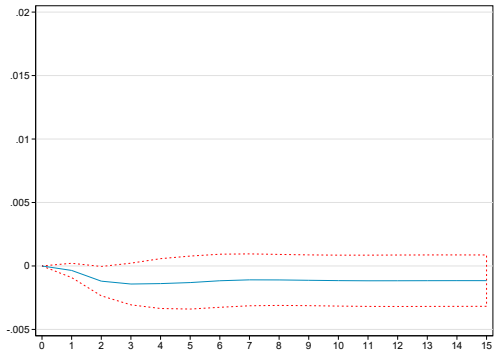
Figure 8: COIRF estimates for the whole period (1996m3 - 2015m10)

Note: This figure reports the Cumulative Orthogonalized Impulse Response Functions (COIRFs) of a one percent shock in all endogenous variables in the system and its impact on aggregate investment. We report the point estimates with one asymptotic standard error band for each impulse response over the horizon of fifteen months. Graphics are in the following order: $\Delta INVESTMENT$, $\Delta CREDIT - GDP$, ΔRIR, ΔRER, $\Delta IBOV$, and EFPI-GDP.

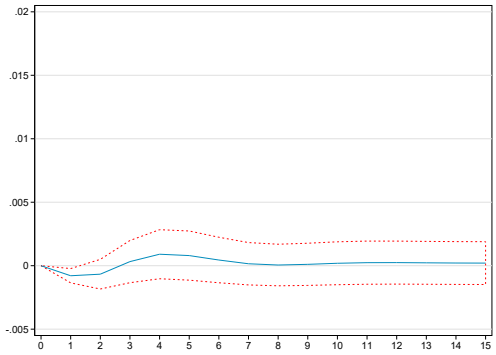
(a) Response of Investment to Investment



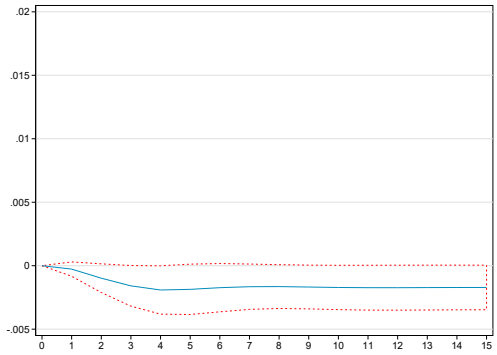
(b) Response of Investment to Credit-GDP ratio



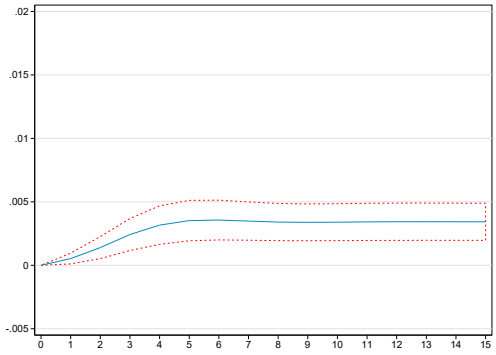
(c) Response of Investment to Real Interest Rate



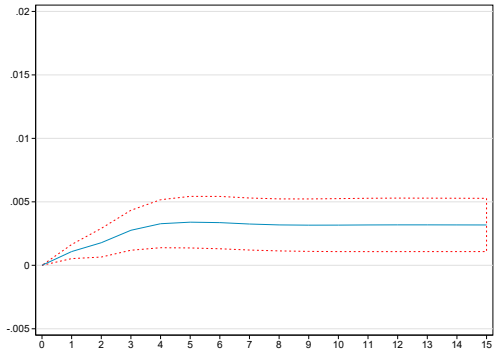
(d) Response of Investment to Real Exchange Rate



(e) Response of Investment to Ibovespa



(f) Response of Investment to Equity Foreign Capital Investment/GDP



Source: Authors' elaboration.

with respect to shocks to credit as a share of GDP, real interest rates and real exchange rates are not statistically different from zero.

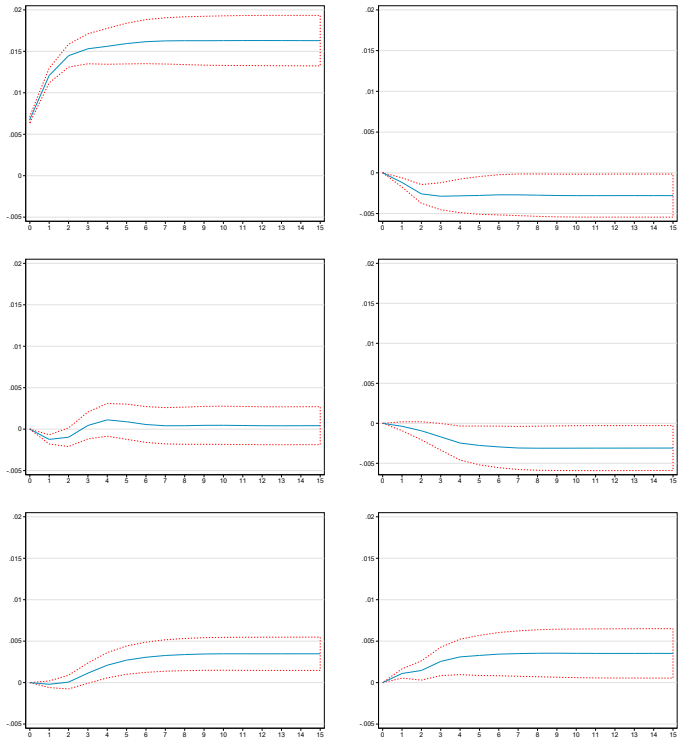
When looking at the sub-periods before and after the 2008 crisis (Figure 9), we see very similar dynamics. In fact, the responses we found for the pre-crisis period are nearly equal to the ones for the full period: investment responds strongly to movements of past investments, but marginally to movements of stock returns and foreign equity capital flows. In the post-crisis period, we see that the response to past movements of investment is still strong, but decays more rapidly, and it seems that the supply of credit takes on some non-negligible importance in explaining movements in investment, as now the response of investment to changes in credit as a share of GDP is positive and statistically significant. Though this is good news, because after the crisis investment becomes sensitive to expanded credit, it seems that it came at the expense of the beneficial effect of foreign capital flows, as after the crisis this variable no longer exerts a positive effect on investment growth.

Overall, from the analysis of Impulse Response Functions we can see that aggregate investment shows a positive, high persistence over time, and thus a positive shock to investment tends to have a positive cumulative effect on its future values, confirming the irreversibility of physical investment hypothesis. Three other variables play a role in determining investment growth: stock market returns, foreign portfolio equity capital flows and supply of credit, but the effect of the later variables is fairly marginal, and are shown to affect investment in different time periods in which the economy was facing different circumstances. In general, all determinants of investment are intimately linked with the neoclassical model, as stock market returns proxy for future expected profits, whereas foreign capitals and supply of credit affect investment through the cost of capital channel. Therefore, to some extent we find some features of the neoclassical model of investment reflected in our estimates.

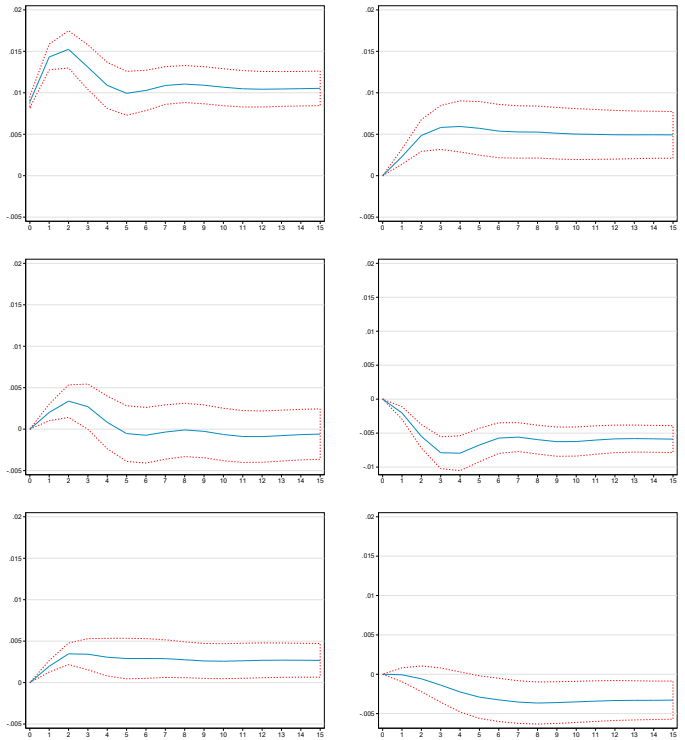
Figure 9: COIRF estimates, before and after the 2008 financial crisis

Note: This figure reports the Cumulative Orthogonalized Impulse Response Functions (COIRFs) of a one percent shock in all endogenous variables in the system and its impact on aggregate investment. We report the point estimates with one asymptotic standard error band for each impulse response over the horizon of fifteen months. Graphics are in the following order for both panels (a) and (b): $\Delta INVESTMENT$, $\Delta CREDIT - GDP$, ΔRIR, ΔRER, $\Delta IBOV$, and EFPI-GDP.

(a) Before the 2008 Financial Crisis (1996m3 - 2008m8)



(b) After the 2008 Financial Crisis (2008m12 - 2015m10)



Source: Authors' elaboration.

2.3.4 Forecast Variance Error Decomposition (FEVDs)

According to Stock and Watson (2001), the forecast variance error decomposition (FEVD) is the percentage of the variance of the error made in forecasting a variable (e.g. aggregate investment) due to a shock in one of the endogenous variables (e.g., EFPI). This variance error is decomposed in a given horizon, usually compatible with the IRF analysis. Since the shocks stabilize after the first months, we compute the FEVDs in the twelve months horizon.

Table 15 shows the results of the variance decomposition of forecast errors. Just as suggested by the estimated parameters of the VARX and by the COIRFs, innovations in real aggregate investment are responsible for about 84% of fluctuations in future investment, revealing a large persistence of physical investment. Meanwhile, other factors account for only 6.8% of the aggregate investment future movements, highlighted by EFPI-GDP (approximately 2.0%), Ibovespa (1.9%), and real interest rate (around 1.5%). We can conclude that, even though EFPI is a statistically significant variable in the system, just a small fraction (around 2.0%) of the forecast error variances of investment are accounted for by innovations in this variable (even so, it is the highest FEVD among all other endogenous variables in the system).

There are differences before and after the crisis, though. Aggregate investment growth shows less persistence after the 2008 financial crisis (Table 15). In this period, innovations to other variables in the system account for 32.7% of the error variance in the investment equation. We can observe an increasing importance of RER in explaining movements in investment, starting 3 months ahead (ending up accounting for 13.8% of the variance of forecast errors). As we observe from the IRFs analysis, a positive shock to RER lead to a negative response in investment, specially after the crisis.

Respective to the role of EFPI on investment, the FEVD analysis suggest that it is more important in explaining variations in investment in the pre-crisis period (3.0% of the error variance 15-steps ahead) rather than in the after period (1.3% 15-steps ahead). These results are consistent with those we find in the IRF analysis and suggest that EFPI play a more important role in explaining future variations in aggregate investment before the 2008 financial crisis, when government interventions in the economy were low.

Table 15: Point estimates of the Forecast Error Variance Decompositions, before and after the 2008 financial crisis

Note: This table presents the fraction of the s-step forecast-error variance of variable i that can be attributed to the jth orthogonalized innovation. We show the FEVDs for the full period and also for out splited samples (before and after the 2008 financial crisis).

Period	INVESTMENT	CRED-GDP	RIR	RER	IBOV	EFPI-GDP
			Full period			
1	100	0.000	0.000	0.000	0.000	0.000
2	98.22	0.096	0.488	0.055	0.222	0.918
3	96.54	0.610	0.467	0.426	0.748	1.209
4	94.21	0.630	1.150	0.673	1.481	1.859
5	93.50	0.613	1.351	0.728	1.820	1.991
6	93.44	0.612	1.347	0.723	1.890	1.984
7	93.34	0.625	1.428	0.734	1.889	1.982
8	93.27	0.628	1.486	0.737	1.891	1.988
9	93.26	0.627	1.491	0.736	1.893	1.989
10	93.26	0.628	1.493	0.737	1.893	1.989
11	93.25	0.628	1.498	0.738	1.893	1.989
12	93.25	0.628	1.499	0.738	1.893	1.989
			Before crisis - 1996m3 to 2008m8			
1	100	0.000	0.000	0.000	0.000	0.000
2	94.50	1.781	1.940	0.174	0.051	1.550
3	92.08	3.885	1.835	0.523	0.125	1.553
4	87.11	3.732	3.896	1.114	1.387	2.756
5	85.04	3.642	4.282	1.721	2.312	2.999
6	84.57	3.619	4.308	1.803	2.691	3.011
7	84.32	3.613	4.408	1.831	2.807	3.026
8	84.24	3.609	4.427	1.848	2.852	3.027
9	84.22	3.610	4.426	1.849	2.864	3.028
10	84.22	3.612	4.428	1.848	2.868	3.028
11	84.21	3.612	4.428	1.848	2.869	3.028
12	84.21	3.612	4.429	1.848	2.869	3.028
			After crisis - 2008m12 to 2015m10			
1	100	0.000	0.000	0.000	0.000	0.000
2	86.19	4.250	3.229	3.261	3.071	0.004
3	73.17	7.878	3.963	10.71	4.104	0.175
4	70.39	7.854	3.934	13.47	3.789	0.561
5	69.38	7.436	5.806	12.75	3.658	0.970
6	68.04	7.265	6.695	13.24	3.575	1.191
7	67.56	7.271	6.664	13.71	3.546	1.249
8	67.53	7.252	6.724	13.67	3.534	1.288
9	67.44	7.242	6.751	13.73	3.540	1.295
10	67.39	7.245	6.760	13.77	3.547	1.295
11	67.33	7.242	6.833	13.75	3.544	1.298
12	67.29	7.236	6.860	13.76	3.542	1.302

Source: Authors' elaboration.

2.3.5 Sensitivity analysis: Does the variables ordering matter?

One of the main critiques to the identification in VAR models using the Choleski decomposition of the residual covariance matrix is that IRF results are ordering-dependent (KILLIAN, 2013). The more correlated the residuals are, the more sensible responses can be to different orders of variables in the system. To get things even more complicated, the ordering of variables in impulse response analysis cannot be determined with statistical methods, and thus has to be specified by the analyst (LÜTKEPOHL, 2010). As previously stated, our baseline variable ordering in the VARX system tries to follow economic theory and relies on economic grounds, but ordering is not a trivial task and sometimes it can be arbitrary[42]. Since this is an important issue related to the interpretation of the IRFs, we then proceed to checking if the results are robust to ordering (a robustness check that is also suggested by Sims, 1981).

In this exercise, we make two assumptions:

- Assumption 1: aggregate investment remains at the top of the vector of the K = 6 endogenous variables, following the idea that physical investment has a higher degree of rigidity and thus tend to react to shocks in other variables with lags;

- Assumption 2: foreign equity portfolio investment remains at the bottom of the vector of endogenous variables, following the idea that it typically refers to foreigners transactions in securities/assets that are very liquid, i.e., these securities can be bought and sold easily and fast, and thus it can react immediately to shocks in other variables.

All other endogenous variables ($\Delta CRED - GDP$, ΔRIR, ΔRER, and $\Delta IBOV$), are then tested in every possible combination in the model, in order to check if their order affects our results. Since we have 4 variables permuting without repetitions, we have 4! = 4 x 3 x 2 x 1 = 24 different combinations.

As shown in Figure 10, the results of this tests show that the cumulative orthogonalized response of investment to a shock in foreign equity portfolio investment in the 15 months horizon

[42] In our baseline variable ordering, for example, it is assumed that real interest rate is not contemporaneously affected by Ibovespa returns, meanwhile Ibovespa returns are affected by contemporaneous shocks in the real interest rate. One could argue that it is reasonable to expect that Ibovespa returns do have an immediate effect on real interest rate, and that is why checking different variable ordering becomes important.

is positive for 100% of the combinations (24/24). Considering the 68% confidence interval for the IRFs, we can reject the null hypothesis that the point estimate is equal to zero 15 months after the shock in EFPI in 54.2% (13/24) of the times. Albeit the CI do not affirm there is a statistically significant impact of EFPI in aggregate investment in all cases, the point estimate is always positive, suggesting that our results are not driven by a specific ordering of variables in the Choleski lower triangular decomposition of the residuals. Instead, these simulations reaffirm the magnitude of the impact: a one unit increase in EFPI-GDP ratio (1 percentage point) tend to have an accumulated impact on the real aggregate investment growth 15 months ahead of 0.3%, holding everything constant.

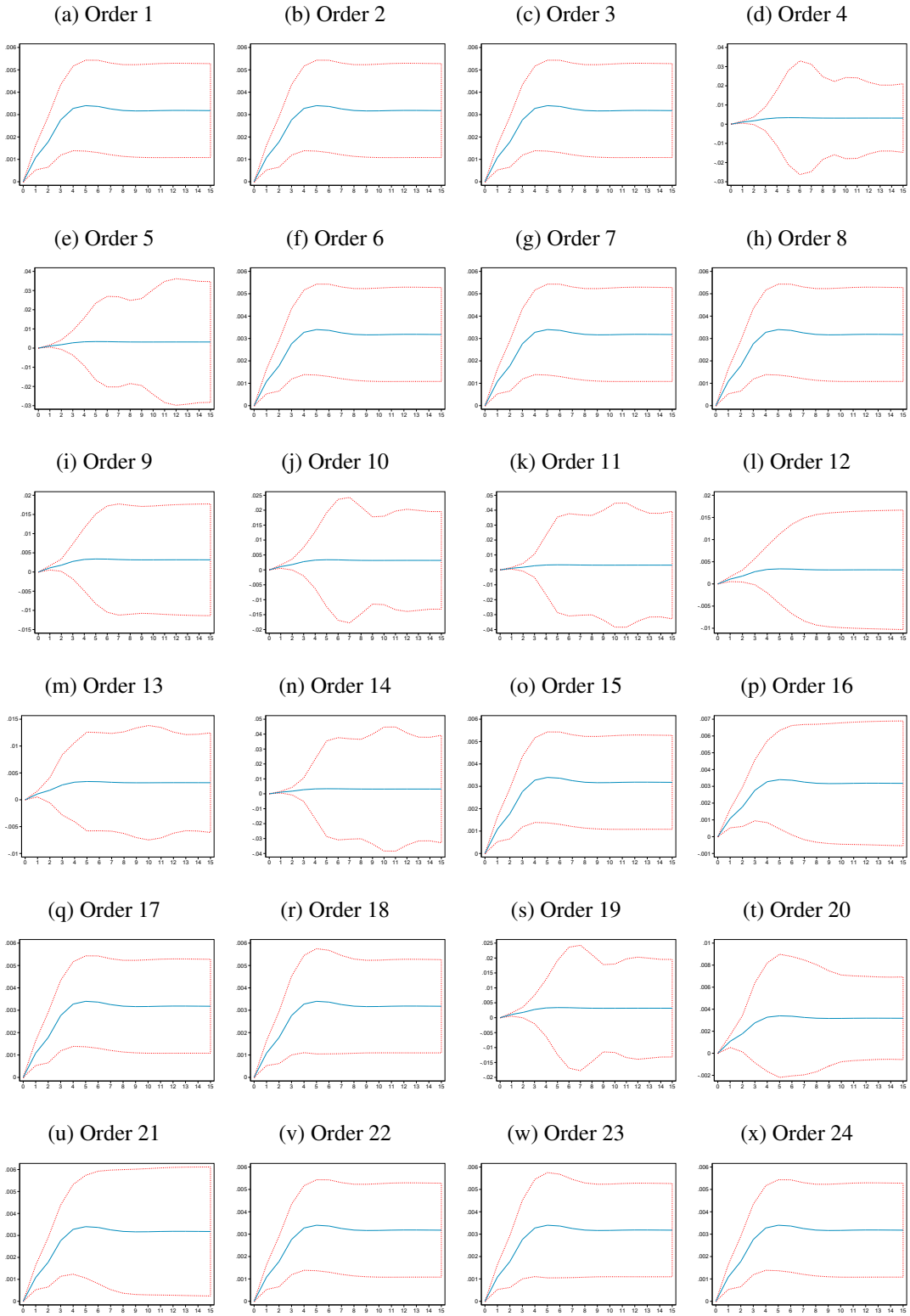
We also replicate this ordering robustness check for before and after the structural break imposed by the 2008 financial crisis (non-reported because of space constraints). We find that the a shock to EFPI caused an average response in aggregate investment growth 15 months ahead of 0.3% (considering the 24 different model specifications). Before the crisis, it averages 0.4%, while after the crisis the average COIRF drops to -0.3%. This contrast between pre and post-interventionism suggest that indeed the role of EFPI on aggregate investment change significantly before and after the crisis, despite of the variables ordering.

Finally, we also calculate the FEVDs for these different variables ordering. On average, the fraction of the forecast errors due to shocks in EFPI is 1.2% in the full period, 1.5% before the 2008 financial crisis, and 0.7% after the crisis, considering the same 24 possible combinations. Taken together, this exercises confirm that EFPI plays a moderated role in explaining future movements in aggregate investment, even before the 2008 financial crisis, where its effects are larger. We can conclude than our results are qualitatively and quantitatively similar if we consider different order variables in the Choleski decomposition of the covariance matrix of the residuals.

2.4 Discussion

In this subsection, we discuss two factors directly related to our results: whether the role of EFPI in aggregate investment changed after the 2008 great financial crisis and how the new Brazilian fiscal reality can amplify the importance of foreign investment flows in the next decades.

Figure 10: Robustness check: COIRFs with alternative variables ordering - Full Period (1996m3 - 2015m10)



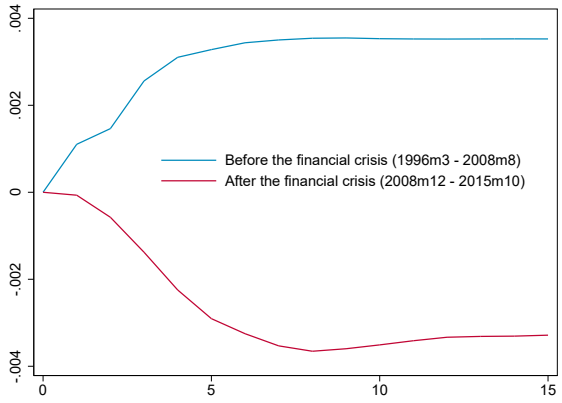
Source: Authors' elaboration.

2.4.1 Did the role of EFPI in Investment change in the recent period?

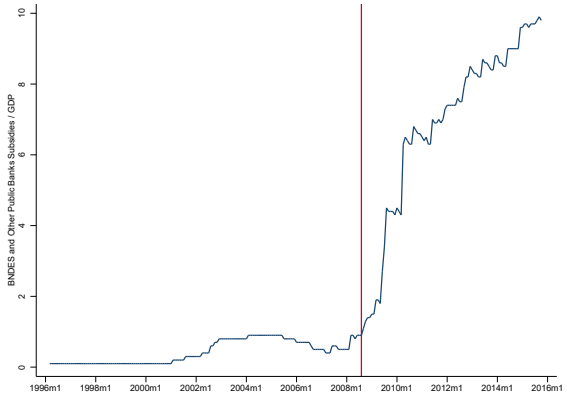
The results from our previous analysis show that the role of foreign capitals in financing investment growth suffered a setback after the 2008 financial crisis. Figure 11 shows the COIRFs of a shock in EFPI to Investment (A) and the Transfers from Brazilian National Treasury to Public Banks as a share of GDP (B). As one could notice, the role of foreign capital inflows in affecting aggregate investment indeed seem to have changed dramatically after the 2008 financial crisis.

Figure 11: The effect of the EFPI on Investment and Structural change in subsidized credit, before and after the 2008 financial crisis

(a) Response of Investment do a shock in EFPI - Before and After the crisis



(b) Transfers from Brazilian National Treasury to Public Banks (% of GDP)



Source: Authors' elaboration.

A fundamental question remains: why has the virtuous cycle between foreign capitals and investment ceased? We offer two arguments. First, the financial crisis has probably interrupted the process of equity market integration in emerging markets. Equity market integration takes time to kick in, and the process might suffer reversals in times of instability (CARRIERI; ERRUNZA; HOGAN, 2007). Indeed, as pointed by Fratzscher (2012), the global crisis caused substantial reallocation of capital from emerging back to developed countries, as institutional investors chased safe-haven investments in low-risk countries. Hence, the crisis might have reduced the flow of foreign capital to the Brazilian equity market, and such capital reallocation seems a natural suspect to explain why the effect of foreign capitals on investment has ceased.

Second, the Brazilian government has shifted economic policy in two main grounds after the crisis. First, by increasing subsidies to large firms, second by enacting several capital control

measures between years 2009 and 2011, motivated by concerns related to financial instability and exchange rate overheating, precisely because foreign capitals resumed to fly in at pre-crisis levels (or even at higher levels) as global markets settled and investors' confidence was relatively restored. Hence, the very central bank might have armed the trap which prevented foreign capitals to continue financing investment, by taking active measures to reduce foreign capital flows. Indeed, putting together these two arguments, an important and current debate in Brazil is whether the increased state intervention in the economy contributed to the recessionary period that started in 2014m3 and is still in course, which clearly has its roots in the preceding years.

The so called "New Economic Matrix" was introduced by President Dilma Roussef through economic incentives for selected industries, without the presence of clear criteria for the granting of benefits such as tax relief. One of the facets of this intervention is the rapid growth of transfers from the National Treasury to Public Banks, which rose from 0.9% of GDP in 2008m8 to 9.8% of GDP in 2016m10 (see Figure 11 - b). The destination of these resources is also often contested, either because many of the benefited companies are large - which can get credit through private banks or the issue of securities in the capital market - or because there is evidence that these loans are guided by political motivation and their availability coincide with electoral periods (CARVALHO, 2014). For many, the Brazilian model of capitalism converged very close to the "crony capitalism" as described by Zingales (2014), where specific groups see more advantage in investing in lobbying than in expanding their productive activity. As we can see, this recent interventionism may have affected the determinants of aggregate investment, specially those related with private investment.

In fact, as theorized by Stulz (2005), foreign equity capital is beneficial to emerging economies only if contracting is efficient too and agency costs are mitigated. Such shift towards interventionist policies that occurred in Brazil increases the discretion of the local government in extracting benefits from firms, via both official and unofficial channels, the so-called agency cost of state ruler discretion. A vivid evidence of such agency costs which arise from inefficient contracting and excessive government interventionism was recently witnessed by the whole world as the corruption scandal in giant oil company Petrobras (which has a mixed public-private ownership structure) unfolded, sparkling a strong response from the public against corruption and excessive state intervention in the Brazilian economy. To make better use of foreign capitals so as to allow an increased effect on investment, Brazil has to improve on the quality of institutions and on the efficiency of

contracting and corporate governance standards as well, and to achieve so government and firms have to work together.

2.4.2 Looking to the future: the new Brazilian fiscal reality and the importance of foreign capital inflows

Finally, there are two actual factors related with our paper that deserve attention. First, the expansionary monetary policy adopted by Central Banks of advanced countries after the 2008 financial crisis renewed the debate over policy options in emerging markets to deal with capital flows (MAGUD; REINHART; VESPERONI, 2014). As of today, the expansionary monetary policies – that keep active even though we are almost ten years from the crisis – seem to be a likely scenario for the next years. These low interest rates in the central financial markets can boost future capital flows to emerging markets, inspired not only by higher returns but also by the fact that for the first time in decades developing countries faced a global financial crisis without feeling heavily its effects.

A second important topic relating to capital flows and its importance in the foreseeing future to emerging markets is the infrastructure gap and the lack of funds to finance investments in areas such as highways, ports and airports. In the Brazilian case, it is estimated that the inefficiencies due to inadequate infrastructure implies a cost of approximately 10-15% of the countries' GDP (CREDIT SUISSE, 2013). Recent study of Garcia et al. (2015) shows that Brazil have inferior overall infrastructure quality relative to almost all its export competitors. From 1980 (around 5%) to 2013 (around 2% of GDP), total infrastructure investment (public + private) shrunk by half of its size when compared to countries' GDP. Recent announcements of a new concession project from the Federal Government bring expectation on attracting private companies – most of them potentially financed by foreign capital – to invest in areas such as transport, energy, and telecommunications. Summarizing, foreign equity capital can be a powerful allie not only for Brazil, but for other emerging markets trying to access funding from the international capital markets to finance its investment needs.

2.5 Conclusion

A recurring question of academics and policy-makers is whether foreign capital flows to emerging markets have real effects on investment and GDP growth. Although we can cite several theoretical benefits from the increasing financial integration – e.g., improving risk sharing, alleviating financial constraints, increasing market liquidity, and forcing better corporate governance practices –, there is still little evidence of its effects in emerging economies. Empirical analysis usually focus on liberalization events (HENRY, 2000; LAEVEN, 2002; BEKAERT; HARVEY; LUNDBLAD, 2005; CHARI; HENRY, 2008), and not on financial integration as a slowly and time-varying process. We try to fill this gap by analyzing the effects of foreign portfolio investments in a major emerging market such as Brazil, where limited funding availability and the high cost of capital limits the development of private investment.

We start by proposing a monthly estimate of the Brazilian gross domestic capital formation that do not affect the original properties of the country's quarterly domestic investment, released by IBGE. This high-frequency local investment data allows us to significantly increase the number of observations, and thus include a larger number of parameters to be included in the VARX modeling. We use this framework to estimate the response of domestic investment to an impulse in several endogenous variables, emphasizing the marginal impact of shocks to foreign portfolio flows.

Overall, our evidence suggests that equity foreign portfolio investment has a statistically significant impact on investment in Brazil, but this effect seems to be economically modest. From 1996 to 2015, our estimates suggest that a one percentage point positive shock in EFPI-GDP leads to a 0.4% increase in domestic investment growth, ceteris paribus. Interestingly, the Brazilian experience shows that the relation between foreign portfolio investment and gross capital formation is time variant and possibly conditional to the degree in which government intervenes in the economy. Following the 2008 financial crisis, transfers from Brazilian National Treasury to State-owned Banks increased from 0.9% of GDP in 2008m8 to 9.8% of GDP in 2015m10. Meanwhile, despite net foreign capital flows remains positive (averaging 0.7% of GDP after the 2008 financial crisis, but decelerating to 0.3% at the end of 2015), responses of domestic investment to shocks in EFPI-GDP became a lot weaker.

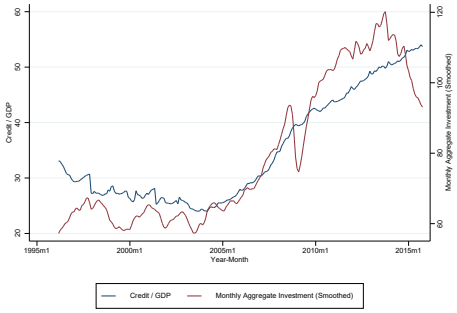
This duality between before and after the 2008 crisis is consistent with the idea that exces-

sive government intervention in credit markets can neutralize the impact of EFPI on investment. Recent domestic investment behavior suggests that growing fiscal deficits and diving business confidence may dominate the positive effect of foreign capital inflows. Looking forward, Brazilian new fiscal reality and the necessity of expanding infrastructure projects in areas such as energy, transportation, and telecommunications may lead to an increasing importance of foreign direct and portfolio investments. In this sense, institutional factors and fiscal deficits sustainability seems to be a still open challenge for the country to attract foreigners investors and to shape the effects of foreign capital on investment and economic growth.

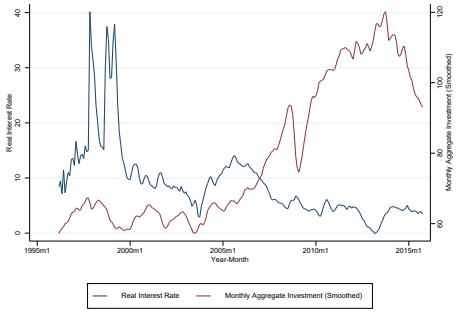
Appendix A: Graphic of Variables - Expressed in Levels - Full Period

Figure 12: Graphic of variables expressed in levels, together with the monthly investment, 1996m3 to 2015m10

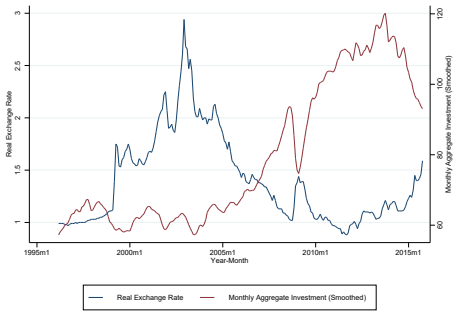
(a) Investment vs Credit-to-GDP



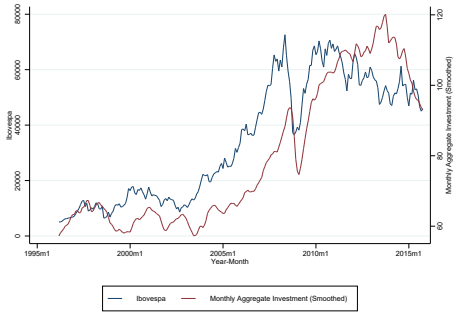
(b) Investment vs Real Interest Rate



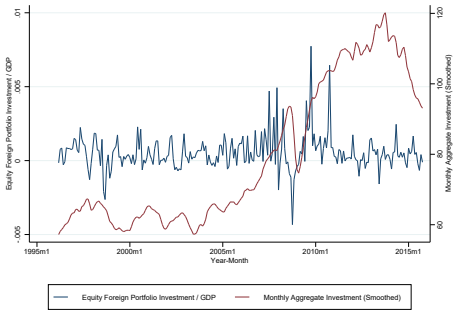
(c) Investment vs Real Exchange Rate



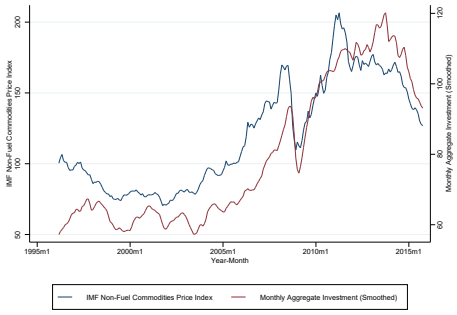
(d) Investment vs Ibovespa



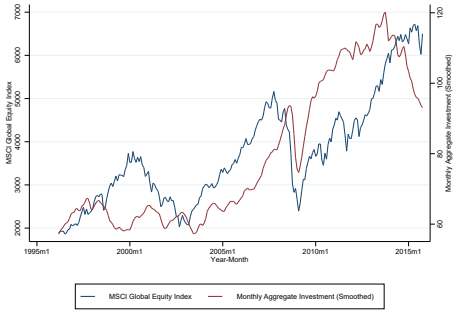
(e) Investment vs Net Foreign Portfolio Flows



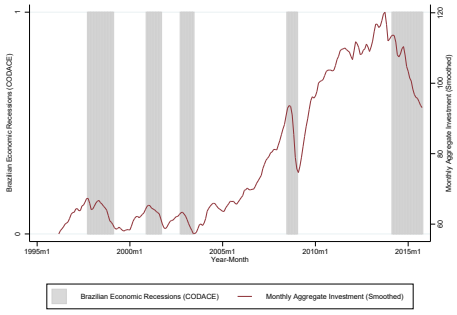
(f) Investment vs Commodity Prices (Ex-fuel)



(g) Investment vs MSCI World



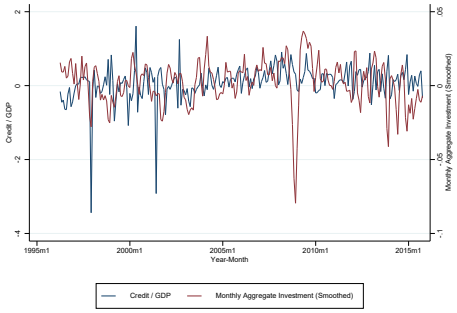
(h) Investment vs Brazilian Economic Recessions



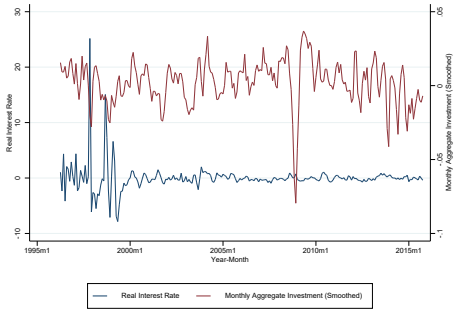
Appendix B: Graphic of Stationary Variables - Full Period

Figure 13: Graphic of stationary variables, together with the monthly investment, 1996m3 to 2015m10

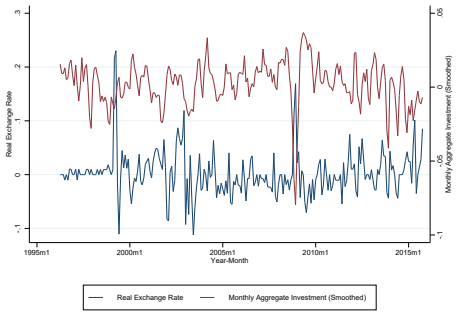
(a) Investment vs Credit-to-GDP



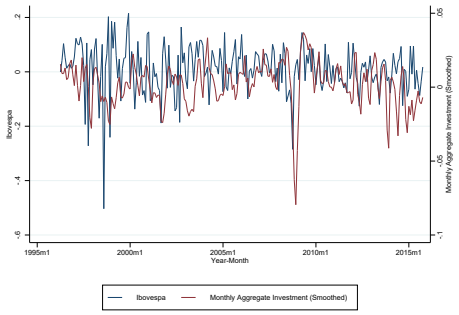
(b) Investment vs Real Interest Rate



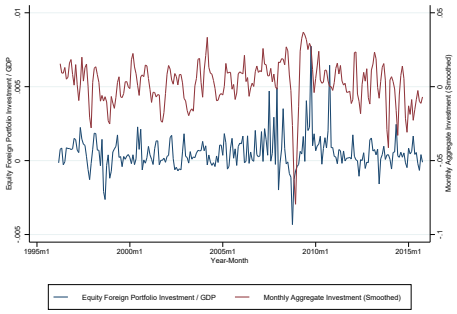
(c) Investment vs Real Exchange Rate



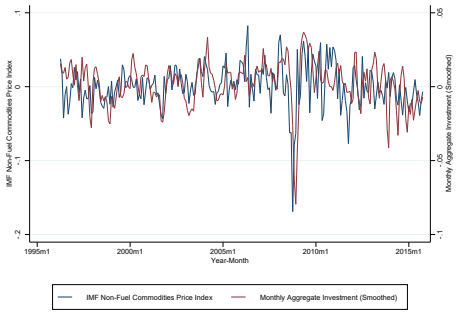
(d) Investment vs Ibovespa



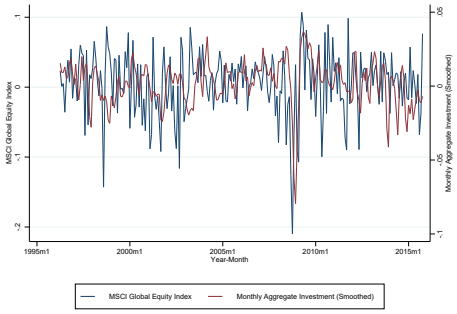
(e) Investment vs Net Foreign Portfolio Flows



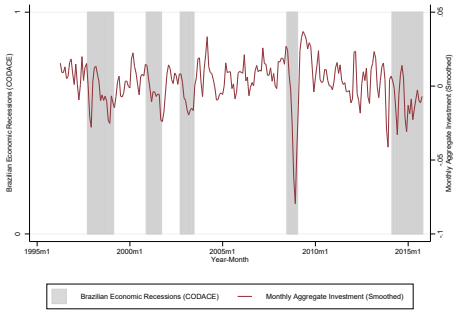
(f) Investment vs Commodity Prices (Ex-fuel)



(g) Investment vs MSCI World



(h) Investment vs Brazilian Economic Recessions



3 Financial constraints, collateral prices, and corporate investment: evidence from Brazil

Abstract

Corporate finance theory points out collateral as a contract instrument that reduces asymmetric information problems and increases value for debtors in default states. During credit expansions, high collateral prices could increase borrowing capacity of firms, especially for those that were financially constrained before the boom period. In this paper, we exploit a real estate prices boom during the 2000s in Brazil to study the role of collateral on corporate financing and investment. Our results suggest that the credit boom of the second half of the 2000s alleviated financial constraints in Brazil, especially for small, less tangible publicly traded firms, ending up to increase corporate investment (weak evidence) and long term debt financing (strong evidence).

Keywords: Financial Constraints; Credit Multiplier; Collateral; Corporate Investment.
JEL Codes: G30, G31, G32.

3.1 Introduction

An important topic in finance research is whether firm-level financial constraints shape corporate investment and how it interacts with the business cycle (ERSAHIN; IRANI, 2015). Evidence from the global financial crisis of 2007-2008 shows that the credit channel affects firms according to the extent that they are exposed to the external shock in the credit market (ALMEIDA et al., 2012), and that constrained firms are likely to end up cutting investments, selling assets, and bypassing investment opportunities (CAMPELLO; GRAHAM; HARVEY, 2010). Financial constrained firms can also impose negative externalities on its industry peers, thus amplifying economic downturns (CARVALHO, 2015). However, little attention has been given to the opposite side of the story: how does constrained firms' investment behaves when access to credit becomes easier in a poorly developed credit market, and how does it interact with economic expansions? These are questions that, to the best of our knowledge, are not well answered in the literature.

One important factor that can limit corporate borrowing when firms have imperfect access to external financing is the value of collateral assets (pledgeable assets). According to Almeida and Campello (2007) model, financially constrained firms could invest below the optimal level because the value of pledgeable assets (such as cash, receivables, inventory, and net property, plant and equipment - PPENT) does not support enough borrowing. During credit booms, the value of collateral assets (such as real estate) can increase significantly, allowing firms to borrow a larger amount for further investment in pledgeable assets. This multiplier effect can integrate credit and economic cycles, amplifying both firm financing and investment growth.

It is important to note that this theoretical relation between tangibility - proxy for firms' collateral assets - and firm's investment is mediated by the ability of firms to access external finance, i.e., it should be expected to be meaningful only to financially constrained firms. If a firm is ex-ante unconstrained, we should expect no strong response of investment because these firms could easily access equity and loan markets even before the structural change in the economy. Changes in the supply of credit are thus more likely to affect small, cash-flow dependent firms, whose investment is potentially limited by the inability of accessing external finance.

In this paper, we focus on an ex-ante incipient credit market such as Brazil, where most firms were likely to face severe credit constraints before the credit deepening process verified on

the late 2000's (ARAUJO; FERREIRA; FUNCHAL, 2012). Because of the country's institutional background, collateral assets assume an important role in shaping bank lending, notably because of information asymmetry problems that arise from low debtors rights (see, for example, STIGLITZ; WEISS, 1981; HART; MOORE, 1994; KIYOTAKI; MOORE, 1997). Lower collateral values, such as property, plant, and equipment - pledgeable assets - can thus limit corporate borrowing, while the credit expansion observed in the country alleviated some of the market frictions and lead the economy to an "above-trend" growth (CARVALHO et al., 2015).

Our identification strategy relies on the enactment of the Law 11.101/2005, the "New Bankruptcy Law", which increased debtor rights and firm access to external finance (ARAUJO; FERREIRA; FUNCHAL, 2012). This regulatory change allows us to use a Difference-in-Difference (DID) methodology to estimate the impact of the law on different categories of firms. Specifically, we use this exogenous source of variation in collateral prices to test the theoretical argument exposed by Almeida and Campello (2007), which, intuitively, predicts that when firms have imperfect access to credit (financial constraints), changes in firm's' ability to obtain external finance may also increase investment. One such source of borrowing capacity is the value of collateral - commonly proxied in the literature by assets tangibility. By mitigating contractibility problems, we should expect a larger effect of collateral prices on external finance access on countries where debtors are not well protected, such as Brazil. Debt collateral is one of the most common enforcement instruments used around the world. In a state of nature where the borrowing firm defaults, lender collateral gives the ability to the lender to seize tangible assets belonging to the debtor (ARAUJO; FERREIRA; FUNCHAL, 2012).

To sort firms on financial constrained and unconstrained groups, we consider the recent Farre-Mensa and Ljungqvist (2016) critique. While some classical measures of financial constraints are still used in the literature of financial constraints (payout ratio, total assets, bond ratings, Kaplan and Zingales' and other indexes, etc.), this paper puts serious doubts on what these variables really measure. The authors find that these measures usually fail in classifying firms that are plausibly constrained, both in debt markets (using natural experiments of 43 staggered increases in corporate income taxes triggered by individual US States) and equity markets (analyzing "equity recycling[43]" on firms). For each of five standard constraints measures, they find that the average "constrained"

[43] Grullon et al. (2011) define equity recycling as the tendency of firms to pay out and raise equity simultaneously.

firm is able to borrow more in response to an increase in state corporate income tax rates and to simultaneously raise equity and increase payouts to shareholders. The authors conclude that these measures could be identifying young and fast-growing firms, rather than capturing financial constraints. Considering their critique, we choose to use firms' total asset as our sorting criteria for financial constraints, which is also consistent with prior studies in the literature (see, for example, KIRCH; PROCIANOY; TERRA, 2014).

Our results show that following the enactment of the New Bankruptcy Law, real estate prices and overall credit in Brazil indeed started to rise faster. For financially constrained firms (bottom 40% of Total Assets, in each year of our sample), we estimate that the low-tangibility group (bottom 40% of its industry fixed assets / total assets) increased investment around 3.0 percentage points (p.p.) more than the high-tangibility group (top 40% of its industry fixed assets / total assets), comparing means before (2002-2004) and after (2005-2008) the exogenous change in collateral prices. We show that this effect is even stronger for firm financing: the average firm in the low-tangibility group increased total debt / total assets around 6.6 p.p. more than the average firm belonging to the high-tangibility group, and this difference is statistically significant at 5%. Finally, this leverage effects on low-tangibility firms are focused on and increase on long-term debt (debt maturing in one year or after), since the low-tangibility group increased the share of long-term debt on total debt around 11.6 p.p. then the high-tangibility group, and this is statistically significant at 1%.

Taken together, our results suggest that the credit boom of the second half of the 2000s alleviated financial constraints in Brazil, especially for smaller, less tangible publicly traded firms, ending up to increase long-term financing and boosting corporate investment. Although we find consistent evidence for the multiplier effect exposed in Almeida and Campello (2007), our results suggest a primary role of banks alleviating collateral requirements, potentially because of the higher supply of credit and better growth opportunities in the economy. In this sense, our evidence supports the idea that the relaxing collateral requirements dominated the multiplier effect during the 2000s credit expansion in Brazil.

Our research connects to at least two research topics. Empirical evidence on the determinants of leverage (FRANK; GOYAL, 2009; HARRIS; RAVIV, 1991; MASULIS; TRUEMAN,

1988; TITMAN; WESSELS, 1988) support the idea that firms that produce steady cash flows and have easily redeployable assets that can be pledged as collateral can afford high debt-equity ratios. In contrast, risky firms with little current cash flows and firms with intangible assets (e.g., RandD and advertising) tend to have low leverage. Our evidence suggests that this wedge can be tightened with credit rights reforms that increase collateral values and access to credit.

Our results are also related to policy experiments using a tax rate reduction in favor of small and medium-sized firms. A recent study on French companies also supports the idea that financial constraints decrease in firm size, as capital accumulation grew more in more productive, small firms affected by the reform (BERNINI; GUILLOU; TREIBICH, 2016). Other recent research shows that a within-firm change in financial constraints can also affect export value, as financially constrained firms reduce their export by 35% relative to financially unconstrained ones (SECCHI; TAMAGNI; TOMASI, 2014). Our study suggests that smaller, less tangible firms are also benefited from a legal reform that strengthened debtor rights in a previously opaque credit market, such as Brazil before 2005.

The rest of the paper is described as follow. In section 2 we discuss the role of collateral in corporate finance, and why it matters especially for emerging economies. In section 3 we describe the credit cycle of the 2000's in Brazil and also its reflects on the value of typically used collateral, such as real estate. In section 4, we describe the data and method that we use. Finally, in section 5, we summarize our results and its main implications on both micro and macroeconomic perspectives, and then we conclude the paper in section 6.

3.2 The role of collateral on corporate financing and investment

Academic studies about the relationship between investment decisions and funding constraints became more popular after the paper of Fazzari et al. (1988). In this study, the authors propose that in the case of firms facing funding constraints, investment decisions are related to the availability of internal resources, and not just the availability of profitable investment opportunities. The study shows, from positive and significant coefficients, the relationship between investment, as the dependent variable, and current and lagged cash flow as independent variables. Note the fact that the cash flow displays explanatory power higher than Tobin's Q. The authors end up validating

the hypothesis that the importance of cash flow on the resources invested is higher in companies with low-profit distribution of dividend.

Since the publication of this study, the theme has been consolidated as one of the most important research topics in financial economics, gaining emphasis in academia through publications both in the international (among others, KAPLAN; ZINGALES, 1997; CLEARY, 1999; ALMEIDA; CAMPELLO; WEISBACH, 2004; CLEARY, 2005; CLEARY; POVEL; RAITH, 2007; ERSAHIN; IRANI, 2015) and national literature (see, for example, TERRA, 2003; KIRCH; PROCIANOY; TERRA, 2014).

Our research, however, is marked out in the model presented by Almeida and Campello (2007). Assuming that imperfections and information asymmetries restrict the ability of firms to take resources and, consequently, to invest, factors that minimize these issues would eventually facilitate their access to lending resources, loosening this restriction. Thus, the authors began to consider the tangible assets subject to garnishment, facilitating access to credit, whereas a larger share of tangible assets would be perceived as an increase in available collateral of the company, making potential funding providers more likely to grant credit to the company. In turn, access to credit would allow greater investment condition in new tangible assets, which could allow new funding and so on. Companies without credit restriction, however, would not be or would be little affected by asset tangibility.

3.2.1 Pledgeable income and its link with collateral

A prospective borrower faces a number of choices. One of them refers to the contract's guarantees - it can be unsecured (lender lending "against cash flow") or secured debt (lenders are lending "against assets"). Pledging assets as collateral is an important instrument to mitigate moral hazard problems and thus reduce interest rates. The model of Almeida and Campello (2007) shows that the value of pledgeable assets can bind firm investment for financially constrained firms.

The model stipulated by Almeida and Campello (2007) assumes the idea presented by Hart and Moore (1994) that human capital can not be alienable, which allows deriving the implication that the pledgeable assets of the company are limited, as well cash flow for investments. The economy has two periods, 0 and 1, and the firm needs to access a particular production technology

$f(I)$, which will generate output in time 1. It is assumed a ratio of external financing (B) to make the project viable. As lenders have no control over the return of the project, it is quite common that start to limit their exposure to the firm value in liquidation, which may be associated with the collateral of the loan agreements. Thus, the resource to be captured (B) shall be limited by the tangibility degree (τ) of Investment (I):

$$B \leq \tau I \tag{9}$$

In addition to the funds raised, the company may also have own resources (W) that can be used to finance investment (I). Like this:

$$I \leq W + \tau I \tag{10}$$

The optimal level of investment (I^{FB}), such that $f'\left(I^{FB}\right) = 1$. If no restriction between own resources and funds raised, the company will be satisfying I^{FB}. However, investment is limited $\left(I^{*} < I^{FB}\right)$, when:

$$\tau < \tau^{*}\left(W, I^{FB}\right) = \max\left(1 - \frac{W}{I^{FB}}, 0\right) \tag{11}$$

If the firm's resources are limited, the level of investment is determined by the available budget. Thus, the optimal investment level is given by:

$$I\left(W, \tau\right) = \begin{cases} \dfrac{W}{1-\tau} & \text{if} \quad \tau < \tau^{*}\left(W, I^{FB}\right) \\ \\ I^{FB} & \text{if} \quad \tau \geq \tau^{*}\left(W, I^{FB}\right) \end{cases} \tag{12}$$

The sensitivity of investment to cash flow is given by:

$$\frac{\partial I\left(W, \tau\right)}{\partial W} = \begin{cases} \dfrac{1}{1-\tau} & \text{if} \quad \tau < \tau^{*}\left(W, I^{FB}\right) \\ \\ 0 & \text{if} \quad \tau \geq \tau^{*}\left(W, I^{FB}\right) \end{cases} \tag{13}$$

Where the sensitivity of investment to cash flow is increasing as the degree of tangibility on investment. However, with a high degree of tangibility, the investment can become disconnected

from the company's cash flow, as well as the degree of tangibility will not impact on the sensitivity of the investment in case of no credit restriction companies.

In this sense, Almeida and Campello (2007) present the following proposition:

Proposition 1 (Almeida and Campello (2007))**.** *The cash flow sensitivity of investment,* $\frac{\partial I}{\partial W}$*, bears the following relationship with asset tangibility:*

i) At low levels of tangibility, $\tau < \tau^* \left(W, I^{FB}\right)$*, the firm is financially constrained and* $\frac{\partial I}{\partial W}$ *increases in asset tangibility.*

ii) At high levels of tangibility , $\tau \geq \tau^* \left(W, I^{FB}\right)$*, the firm is financially unconstrained and* $\frac{\partial I}{\partial W}$ *independent of asset tangibility.*

From this proposal, the model estimated by Almeida and Campello (2007) was:

$$\mathsf{Investment}_{i,t} = \beta_1 Q_{i,t} + \beta_2 \mathsf{Cash\ Flow}_{i,t} + \beta_3 \mathsf{Tangibility}_{i,t} + \beta_4 \left(\mathsf{Cash\ Flow} \times \mathsf{Tangibility}\right)_{i,t} + \sum_i \mathsf{firm} + \sum_t \mathsf{year} + \varepsilon_{t,i} \quad (14)$$

Analyzing the behavior of firms during the period 1971-2000, the study found that the investment-cash flow sensitivities are increasing in the tangibility of firms' assets, but only if firms are financially constrained. These results are consistent with the theoretical predictions of the model.

Still analyzing Almeida and Campello's (2007) model, the credit multiplier drives the relationship between tangibility mechanism of assets and sensitivity of investment to cash flow. With the change in asset prices over time, the credit offer ends ranging jointly, given the movements of the collateral value. Thus, the tangibility effect on companies in the investment sensitivity to cash flow being expanded during periods of economic boom, when the valued assets serve as collateral for making capabilities that will serve to further investment in assets. In times of economic recession, with falling asset prices, companies end up having less sizable assets, which generates fewer guarantees for new lines of credit, containing credit for new investments. The study proves that this credit multiplier has first-order effect on companies' investment decisions.

3.2.2 Non-collateral debt financing: lending on cash

In a perfect market, according to the theory proposed by Modigliani and Miller (1958), investment decisions would merely be affected by the investment opportunities of firms, since the resources would be optimally used and all credit demand would be supplied by supply. In this context, collateral assets are not relevant because interest rates would clear the market, even in the presence of bankruptcy risk.

In the real world, however, market incompleteness and asymmetric information can lead the economy to a credit rationing situation (STIGLITZ; WEISS, 1981): adverse selection, moral hazard, and limited enforcement in case of default cause the exclusion of some individuals on the market, even if they are willing to pay the market value of lending.

Asymmetric information about firms' prospects between insiders and outsiders (potential investors) can create a substantial cost differential between internal and external funds (MYERS; MAJLUF, 1984; GREENWALD; STIGLITZ; WEISS, 1984, and so on.). The break-even q value for a new investment project is higher than unity $(1 + \Omega > 1)$, where Ω is the premium necessary to compensate new investors for the risk of investing inadvertently in lemons[44] (FAZZARI et al., 1988). Given this context, this asymmetric information causes "credit rationing" for some borrowers (STIGLITZ; WEISS, 1981).

To minimize this uncertainty, lenders seek to avoid purely loans based on cash flow and not backed by assets, that is when the expectation of receiving the money back is purely based on the assessment that the borrower will be able to generate enough cash flow. In the case of default, the lenders can not repossess (seize) specified assets. In this case, firms can be credit rationed if they do not have enough pledgeable assets to get credit, and Tirole (2010) establishes the following hypothesis, which we will take to our study:

$\mathcal{H}1$: Credit rationing is more binding for firms with less tangible assets or assets that have a lower value in liquidation (TIROLE, 2010).

[44] Some papers in corporate finance adopted the term "lemon", originally proposed by Akerlof (1970), to describe bad type firms.

3.3 Credit cycles and collateral prices in Brazil

In this subsection, we discuss the recent credit expansion observed in Brazil and the behavior of the real estate collateral prices.

3.3.1 Credit evolution in Brazil

With the implementation of the Real Plan, in 1994, and the consequent economic stabilization from the control of inflation, Brazil opened space for the growth of the loan portfolio, which subsequently sustain economic growth cycle early 2000. Initially, however, restructuring programs of the financial system, such as PROER - Incentive Program for Recovery and Strengthening of the National financial system, PROES - Incentive Program for the Reduction of the State Public Sector in Banking and PROEF - Strengthening Program Federal Financial Institutions, eventually leading to a credit retraction, which fell from 34% of GDP in December 1995 to 26% of GDP in December 2002 (MORA, 2014).

As can be seen in Table 16, the credit expansion cycle started with the acceleration of applications with funding of free resources, to the detriment of earmarked resources that little grew during this period from 2002 to 2008 and only from 2009 it began to have significant growth rates. From observation of the division between funds invested in Individual and Corporations related applications with free funds, we can check that, given the optimistic expectations of a recovery in employment and household income, the initial growth happens in the Individuals segment, passing from 6.1% in 2002 to 13% in 2008. Meanwhile, credit to companies remained stagnant until 2005, starting their growth cycle from 2006 and reaching 15.7% of GDP in 2008.

Still based in Table 16, it is possible to observe that growth is sustained by the appetite of private institutions, and more specifically, by national capital banks. These institutions, which had a portfolio of 16.3% to GDP in 2002, increased a portfolio of credit amounted to 25.8% of GDP in 2008. Although public banks have very significant participation in the current context in Brazil, the growth of these occurs in the period following analyzed here, that is from 2009.

Thus, the loan portfolio, which represented 26% of GDP in December 2002, showed broad growth, reaching, according to Central Bank data (Table 16), 40.5% of GDP in December 2008. However, even after this large growth, the credit portfolio in Brazil still represented a low ratio to

Table 16: Relation between credit and GDP: Period 2002-2008

Note: This table presents the evolution of different credit operations in the Brazilian Financial System, as reported by the Brazilian Central Bank (BCB). Earnmarked funds refer to financing regulated by National Monetary Council (CMN) or linked to earmarked resources, mainly destined to housing, rural and infrastructure sectors. Non-earnmarked refers to financing and loans which rates are freely negotiated between financial institutions and borrowers, i.e., market rates. In non-earmarked operations, financial institutions have autonomy to decide loans destination. We also use BCB's classification of outstanding credit according to the type of customer (non-financial corporations - Corporations - and households - Individuals). Finally, we separate banks in public and private, and the latter in national and foreign, depending on the ownership.

Account	**2002**	**2003**	**2004**	**2005**	**2006**	**2007**	**2008**
Non-Earnmarked Resources	16.3	15.0	16.4	18.8	21.0	24.8	28.7
Individuals	6.1	5.9	7.1	8.9	10.0	11.9	13.0
Corporations	10.1	9.1	9.2	9.9	11.0	12.9	15.7
Earmarked resources	9.8	9.6	9.3	9.5	9.9	10.3	11.7
Total Credit	**26.0**	**24.6**	**25.7**	**28.3**	**30.9**	**35.2**	**40.5**
Public banks	9.8	9.8	9.9	10.4	11.3	12.0	14.7
Private Banks	16.3	14.8	15.8	17.8	19.6	22.0	25.8
National	9.7	9.5	10.2	11.5	12.8	15.4	17.3
Foreign	6.5	5.3	5.6	6.3	6.8	7.8	8.5
Total Credit	**26.0**	**24.6**	**25.7**	**28.3**	**30.9**	**35.2**	**40.5**

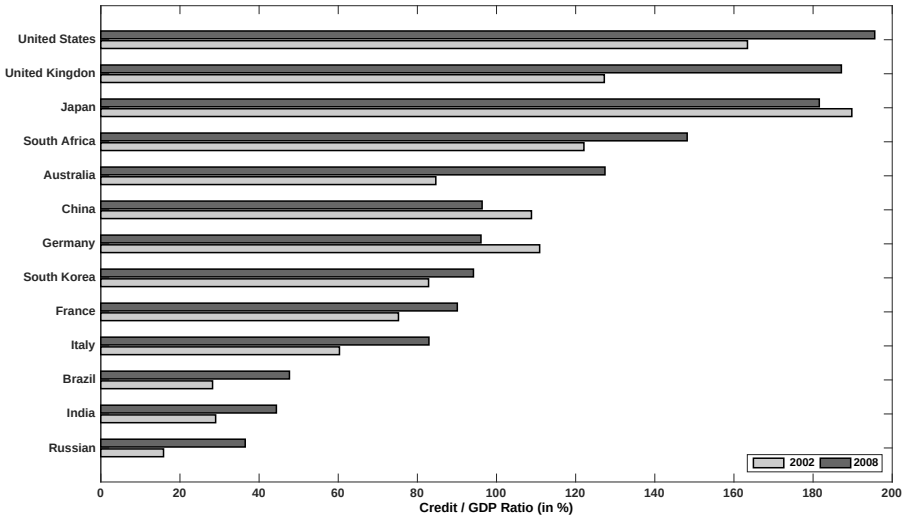
Source: Authors' elaboration.

GDP compared the world's major economies. According to the Financial Development and Structure Dataset (World Bank), presented in Figure 14, the credit / GDP ratio, which was 28.24% in 2002, reached 47.67% in 2008, while reached 195.58% in the United States and even in developing economies such as China and South Africa already was 96.36% and 148.18%, respectively.

Sant'anna, Junior and De Araujo (2009) also point out that the expansion of credit between 2004 and 2008 was accompanied by a major change in market profile both with respect to the extension of deadlines as regarding the fall in interest rates. According to Central Bank data, the average term of the loans, which was 222 calendar days in January 2004, spent 379 days in December 2008, an increase of 70.6% over a period of five years. In turn, the interest rate, which revolved at a level of 45% from 2004 to 2006, came to about 35% between 2007 and 2008.

In Figure 15, we show the Credit-GDP ratio evolution from 1996m3 to 2016m3. Visual analysis suggest that credit growth changed its slope in 2004/2005, following an expansionary

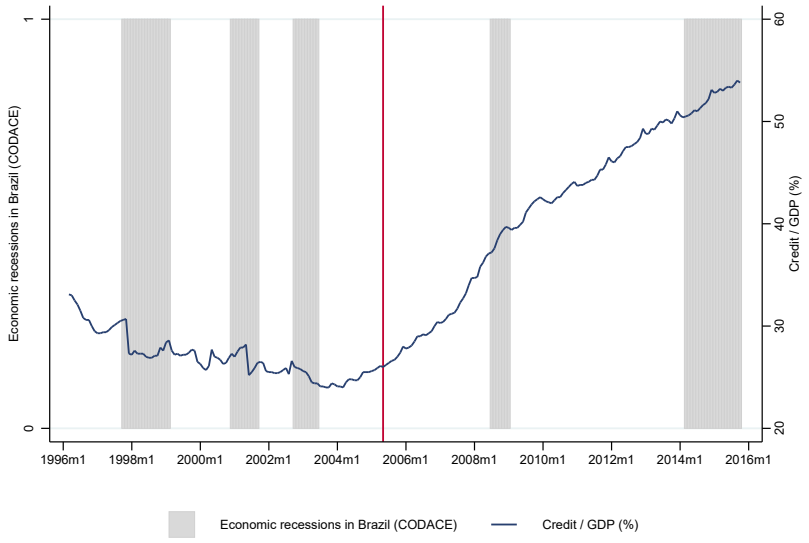
Figure 14: Domestic Credit to Private Sector by Country - 2002 and 2008



Source: Authors' elaboration.

Figure 15: Credit / GDP evolution in Brazil, 1996-2016

Note: This figure shows the evolution of Total Credit / GDP (%) in Brazil (series number 20,622 of the Time Series Management Series, maintained by the Brazilian Central Bank). We plot a vertical red line in 2005m5, when took effect Law n. 11.101/2005, the "New Bankruptcy Law". Shaded areas represent economic recessions in Brazil as dated by the Business Cycles Dating Committee (CODACE).



Source: Authors' elaboration.

economic period of that last 61 months.[45]

3.3.2 Collateral prices in Brazil

To evaluate the time series of pleadgeable assets' price in Brazil, we analyze the Residential Real Estate Collateral Value Index (IVG-R). This index is released by the Brazilian Central Bank and it is calculated with data extracted from the SCR real estate loans to households. This data comprise loans in which the collateral is composed of liens on residential real estate or real estate mortgage. The value of each property backed as collateral is estimated at the time of the granting of credit.

The IVG-R considers the assessed value of the properties pledged as collateral for loans in a given month, from a sample of about 160 thousand households. The collection is made up of properties located in the eleven metropolitan regions considered in calculating the National Index of Consumer Price (IPCA): Belém, Belo Horizonte, Brasília, Curitiba, Fortaleza, Goiânia, Porto Alegre, Recife, Rio, Salvador and São Paulo.

As Figure 16 shows, starting on 2005, we can observe an expansion period on the growth of assets value used as collateral on domestic lending. This acceleration in the growth rate of asset prices occurs exactly in the same period that credit growth accelerated, as seen in the previous session. This increasing value on this type of assets can affect corporate investment through the credit multiplier, as the larger value of pledgeable assets would allow greater access to credit. With more access to credit, companies can execute new investment projects, which generates more pledgeable assets.[46]

3.4 Methodology and Data

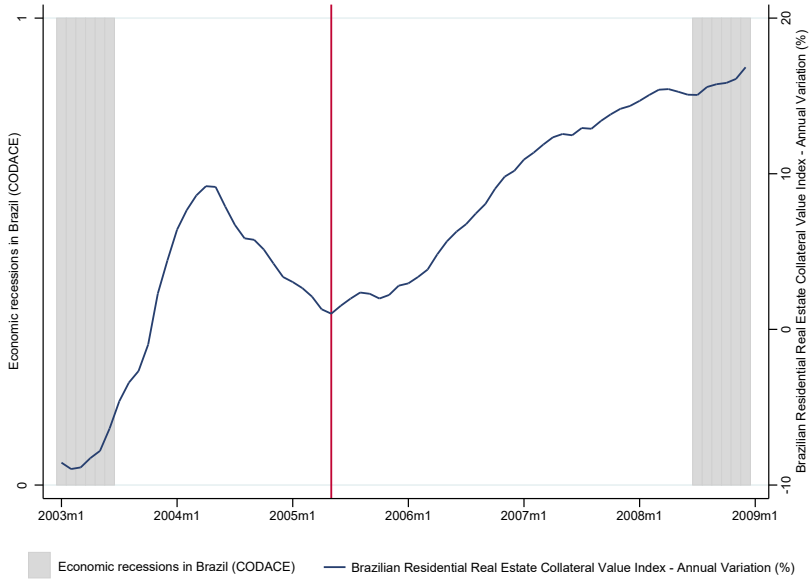
In this subsection, we present the methodology, data and model specification used in the empirical analysis.

[45] This is the largest expansionary period dated in Brazil since data is available (1980). An average expansionary period lasts 28.7 months, while an average recession lasts 15.8 months (CODACE, 2015).

[46] This is the multiplier effect, that could potentially be stronger for firms with more tangible assets. In a macroeconomic point of view, this higher prices of collateral could amplifying corporate investment growth.

Figure 16: Brazilian Residential Real Estate Collateral Value Index, annual growth (%)

Note: This figure shows the evolution of the Brazilian Residential Real Estate Collateral Value Index, calculated by the Brazilian Central Bank (series number 21,340 of the Time Series Management Series, maintained by the Brazilian Central Bank). We plot a vertical red line in 2005m5, when took effect Law n. 11.101/2005, the "New Bankruptcy Law". Shaded areas represent economic recessions in Brazil as dated by the Business Cycles Dating Committee (CODACE).



Source: Authors' elaboration.

3.4.1 Definition of tangibility and investment

Defining asset tangibility is not an easy task. Most countries allow secured debt transactions involving "immovable assets" (eg., land and buildings), while "movable assets" (like machinery and equipment) comprise about half of total fixed assets around the world (ALVAREZ, 2011). We built our baseline measure of tangibility based on the most used proxy in the literature (see, for example, VIG, 2013; CAMPELLO; LARRAIN, 2016), the ratio of total fixed assets (net property, plant and equipment $-$ "movable" $+$ "immovable" assets) to total assets.[47]

$$\text{Tangibility} = \frac{\text{Fixed Assets}}{\text{Total Assets}} \tag{15}$$

Almeida and Campello (2007) use three different measures of tangibility. One is based on a firm-level measure of expected liquidation values according to Berger, Ofek and Swary (1996). Using data from the proceeds of discontinued operations reported by a sample of Compustat firms over

[47] Although Campello and Larrain (2016) also calculate tangibility for U.S. firms using "movable" and "immovable" assets separately, Compustat Global does not comprise this information about the individual components of fixed assets outside the U.S.

the 1984-1993 period, this empirical paper finds that a dollar of book value yields, on average, to a 72 cents in exit value for receivables, 55 cents for inventory, and 54 cents for fixed assets. Almeida and Campello (2007) use these coefficients to estimate a firm-level and time variant "liquidation value":

$$\text{Tangibility}_2 = 0.715 \cdot \text{Receivables} + 0.547 \cdot \text{Inventory} + 0.535 \cdot \text{Capital} \tag{16}$$

The second measure of asset tangibility used by Almeida and Campello (2007) is based on the high-cyclicality of durable goods industry sales. They use a durable/nondurable industry dichotomy to create a measure of assets redeployability, i.e., the extent to which an asset can be (easily or not) sold in the secondary market. The authors include all durable goods industries (except SICs 32 and 38) plus SIC 30, that are assigned with value "1", and "0" otherwise (all other industries - non-durable). Their motivation is that assets of firms operating in non-durables industries are perceived as more liquid by lenders.

In this study, we use in our baseline regressions the Campello and Larrain (2016) measure of tangibility (Equation (15)), since we are interested in the value increasing of real estate assets (immovable assets) and not in other financial items that can be pledged as collateral, such as receivables or inventory. However, we keep in mind that we have different measures of tangibility available in the literature.

Our core investment variable is based on Almeida and Campello (2007) and Kirch, Procianoy and Terra (2014): Investment equal to the ratio between capital expenditures and lagged capital stock (PPENT). We also consider an alternative measure of investment based on Campello and Larrain (2016): the change in fixed assets scaled by lagged fixed assets. Since these variables are highly correlated, we use in our baseline the regressions of the first investment definition.

3.4.2 Data and sample

We start our database collecting Compustat Global Annual Fundamental data for Brazilian publicly traded companies' balance sheet. In order to match this financial data with stock price information, we merge the Compustat dataset with the Center for Research in Security Prices (CRSP)'s daily price securities, using Global Company Key (GVKEY) and fiscal year as matching

variables. The price information we import from the CRSP database is the average price of each firm's stock in a given year. To guarantee that we would lose the minimum number of observations during the matching procedure, we do not impose a minimum number of trading days for each stock in the CRSP database[48].

We also collect aggregate credit information and the Residential Real State Collateral Value Index from Central Bank of Brazil (Central Bank Time Series Management System). Our sample period surrounds the New Bankruptcy Law in Brazil (Law 11,101/2005), going three years forward and three years backward (i.e., 2002-2008). To mitigate the influence of outliers, we winsorize each continum variable in the 5% and 95% trasholds, respectivelly. Following this procedure, we drop firms that did not have at least two year-observations during the period 2002-2008. We also follow Almeida and Campello (2007) and Kirch, Procianoy and Terra (2014) and drop observations (firm-year) that: a) had property, plant and equipment (PPENT) lower than R$5 million, in values of July 2016;[49] b) had asset growth in the year higher than 100%, which typically occurs when firms is involved in mergers and acquisitions (MandA activities); c) were outside the 2 digit SIC range 20-39, that represents manufacturing firms[50].

Our final sample comprises 1,473 firm-year observations (Table 17), divided in "Constrained" and "Unconstrained" subsamples: if a firm in a given year belongs to the top 40% of total assets in its industry (SIC 2 digits), then it is considered unconstrained; and constrained if in the bottom 40%. We use this criterion to separate firms that potentially face low financial frictions (large firms) from those that faces higher financial frictions (small and medium firms), which is consistent with the criteria used in Almeida and Campello (2007) and Kirch, Procianoy and Terra (2014).

Table 17 shows that our sample has a sufficiently large number of high and low-tangibility firms in most of the manufacturing sector. We have a larger fraction of firms from Chemicals and Allied Products (12.2%), Food and Kindred Products (11.8%), and Primary Metal Industries

[48] Even though to calculate the average year price we do not consider missing observations, after the merging process we ended up losing 101 companies (firms for which there was no price or information on the number shares available on the CRSP database). Because our sample is already relatively small, we decided to use the non-merged database in our baseline regressions, using market information such as Tobin's Q only in robustness checks.

[49] We inflate all continuous variables to July 2016 through the national Consumer Price Index - IPCA.

[50] We therefore excludes of the sample firms from the agriculture sector, mining, and also services, such as those firms from the financial industry, typically excluded in empirical corporate finance studies because of its singularities in capital structure and other financial aspects.

Table 17: Stratification of our observations by industry and financial constraint status, 2002-2008

Note: This table reports the number of observations in our final sample, from 2002 to 2008, by industry (SIC 2 digits). The data come from the Compustat Global Database and initially comprises all Brazilian publicly traded firms. Constrained (non-constrained) firms are those in the bottom (top) 40% of firm size (proxied by total assets) in a given industry-year.

Industry	Constrained	Non-constrained	Total	Percent	Cummulative
Food and Kindred Products	77	77	154	11.76	11.76
Tobacco Products	12	0	12	0.72	12.48
Textile Mill Products	62	62	124	8.8	21.28
Apparel, Finished Products from Fabrics and Similar Materials	51	51	102	6.64	27.92
Lumber and Wood Products, Except Furniture	16	11	27	1.68	29.6
Paper and Allied Products	39	39	78	4.64	34.24
Printing, Publishing and Allied Industries	11	7	18	1.2	35.44
Chemicals and Allied Products	91	91	182	12.16	47.6
Petroleum Refining and Related Industries	22	11	33	2.08	49.68
Rubber and Miscellaneous Plastic Products	22	18	40	2.56	52.24
Leather and Leather Products	14	11	25	1.36	53.6
Stone, Clay, Glass, and Concrete Products	29	29	58	3.44	57.04
Primary Metal Industries	79	79	158	11.04	68.08
Fabricated Metal Products	59	59	118	8.24	76.32
Industrial and Commercial Machinery and Computer Equipment	60	60	120	8.64	84.96
Electronic and Other Electrical Equipment and Components	31	31	62	4.16	89.12
Transportation Equipment	63	63	126	8.72	97.84
Measuring, Photographic, Medical, and Optical Goods, and Clocks	14	9	23	1.44	99.28
Miscellaneous Manufacturing Industries	12	0	12	0.72	100
Total	765	708	1,473	100	100

Source: Authors' elaboration.

(11.0%). Industries that had not at least three firms in the sample were dropped from the sample.

3.4.3 Model Specification

To quantify the effect of collateral prices on firm outcomes, we start by estimating a standard difference-in-differences specification using the validity of the law number 11,001/2005, the "Bankruptcy Law", as an exogenous variation in the price of collateral:

$$Y_{is,t} = \alpha_i + \alpha_t + \beta_1 \cdot \mathsf{Post}_t + \beta_2 \cdot \mathsf{High\ Tangibility}_{is} + \beta_3 \cdot (\mathsf{Post}_t \times \mathsf{High\ Tangibility}_{is}) + \varepsilon_{is,t} \quad (17)$$

Where $Y_{is,t}$ represents the outcome of interest (Debt/Total Assets, Long-Term Debt/Total Debt, and Investment/Total Assets), $\mathsf{High\ Tangibility}_{is}$ is a dummy variable that equals one if the firm is in the top 40% of its industry's tangibility (Fixed Assets/Total Assets) distribution, and zero if it is in the bottom 40%; Post_t is a dummy variable that equals one if in the year 2005 or after (this period is determined by the Brazilian New Bankruptcy Law - Law n. 11,101, passed on congress on February 9th, 2005 and took effect three months after) and 0 in years before 2005; to control for firms' time-invariant characteristics we include a full set of firm fixed effects (α_i), as well as year

fixed effects (α_t) to control for time-varying economic shocks. Finally, $\varepsilon_{is,t}$ is the error term, with standard errors clustered at the firm level (PETERSEN, 2009). Our coefficient of interest is β_3, which measures the pre-post difference in the outcome of firms belonging to the high tangibility group, relative to the pre-post difference of firms in the low tangibility group.

Equation 18 estimates the DID controlling for firms and year fixed effects, however, it does not control for other firm and industry characteristics that can simultaneously affect the outcome of interest. To deal with that, we include in Equation 18 a set of control variables, with turns our DID equation similar to Campello and Larrain (2016):

$$Y_{is,t} = \alpha_i + \alpha_t + \beta \cdot (\mathsf{Post}_t \times \mathsf{High\ Tangibility}_{is}) + \gamma X_{is,t} + \varepsilon_{is,t} \tag{18}$$

where $X_{is,t}$ is a vector of firm-level controls that include size, profitability, leverage, and overall tangibility. We use specifications (17) and (18) in the main empirical results of the paper.

3.5 Results

We here show the main results of our empirical analysis. We start by looking at some descriptive statistics, and then we discuss the effects of tangibility on firm financing and investment.

3.5.1 Descriptive statistics

Table 18 presents the descriptive statistics by financial constraint status. Numbers referrers to constrained (Panel A), unconstrained (Panel B) and all firms in the sample (Panel C). As previously stated, the separation of constrained and unconstrained firms is due to within industry total assets (firm size). For this reason, average constrained firm' size (LN Total Assets = 5.15) is lower than the average non-constrained firms' size (LN Total Assets = 7.51), even though firms are in the same industry.

The investment of the two groups has a sensible difference. While the group of constrained firms presents investment of 17% of lagged fixed assets, the group of unconstrained firms features 21% of lagged fixed assets, a difference of 4% of assets, and the group formed by all companies obtained an average of 19% of fixed assets. Likewise, the median of the unconstrained group (0.11) was also lower than the unconstrained group (0.16).

Table 18: Descriptive statistics by financial constraint status and total, 2002-2008

Note: This table reports the descriptive statistics (by financial constraint status) of the firms in our final sample, from 2002 to 2008. The data come from the Compustat Global Database and initially comprises all Brazilian publicly traded firms. Constrained (non-constrained) firms are those in the bottom (top) 40% of firm size (proxied by total assets) in a given industry-year.

Variables	Mean	Median	Std Dev	Minimum	Maximum	# of Obs.
	Panel A: Financial Constrained Firms					
Firm Investment	0.17	0.11	0.21	0.00	2.11	655
Tangibility_1	0.40	0.37	0.21	0.02	0.90	765
Tangibility_2	0.80	0.86	0.18	0.13	1.00	765
Size	5.15	5.08	1.29	2.07	9.41	765
Capital	0.40	0.37	0.21	0.02	0.90	765
Cash	0.08	0.03	0.11	0.00	0.75	765
Δ Cash	0.00	0.00	0.08	-0.79	0.29	765
Cash Flow	-0.03	0.05	0.36	-3.25	0.86	646
Total Debt	0.59	0.25	1.22	0.00	12.35	765
Δ Total Debt	0.15	0.02	1.87	-10.04	39.66	765
Short Term Debt	0.61	0.63	0.31	0.00	1.00	735
Long Term Debt	0.39	0.37	0.31	0.00	1.00	735
	Panel B: Financial Unconstrained Firms					
Firm Investment	0.21	0.16	0.23	0.00	3.50	655
Tangibility_1	0.38	0.38	0.17	0.03	0.87	708
Tangibility_2	0.82	0.86	0.14	0.10	1.00	708
Size	7.51	7.36	1.46	4.33	12.64	708
Capital	0.38	0.38	0.17	0.03	0.87	708
Cash	0.13	0.11	0.11	0.00	0.65	708
Δ Cash	0.01	0.01	0.08	-0.42	0.53	708
Cash Flow	0.08	0.09	0.11	-1.09	0.35	655
Total Debt	0.31	0.3	0.18	0.00	1.68	708
Δ Total Debt	0.03	0.02	0.12	-0.61	0.56	708
Short Term Debt	0.47	0.44	0.26	0.00	1.00	704
Long Term Debt	0.53	0.56	0.26	0.00	1.00	704
	Panel C: Total					
Firm Investment	0.18	0.14	0.19	0.00	3.50	1310
Tangibility_1	0.39	0.38	0.19	0.02	0.87	1473
Tangibility_2	0.81	0.86	0.16	0.10	1.00	1473
Size	6.25	6.23	1.73	2.07	12.64	1473
Capital	0.39	0.38	0.19	0.02	0.90	1473
Cash	0.10	0.07	0.10	0.00	0.75	1473
Δ Cash	0.01	0.00	0.08	-0.79	0.53	1473
Cash Flow	0.05	0.07	0.13	-3.25	0.86	1301
Total Debt	0.30	0.26	0.22	0.00	12.35	1473
Δ Total Debt	0.09	0.02	1.41	-10.04	39.66	1473
Short Term Debt	0.55	0.53	0.30	0.00	1.00	1439
Long Term Debt	0.45	0.47	0.30	0.00	1.00	1439

Source: Authors' elaboration.

Regarding the tangibility, we calculate two measures, namely the relationship between fixed assets and total assets (Tangility$_1$) and the ratio of the sum of cash, accounts receivable, inventories, and fixed capital to total assets (Tangibility$_2$). The results of the two groups were very similar, being that the group of constrained firms (0.40) obtained Tangibility$_1$ slightly higher than the group of unconstrained firms (0.38), while the unrestricted group (0.82) obtained Tangibility$_2$ slightly higher than constrained group (0.80).

With regard to debt, the group of constrained firms showed more debt in relation to total assets, with an index of 0.59 compared to 0.31 of the group of unconstrained firms and 0.45 for the whole group. Interesting to note that in the period, the average change in debt was 0.15 for constrained firms against 0.03 of unconstrained companies, clearly demonstrating that companies of the first group contracted more debt than the second. Similarly, the group of unconstrained firms demonstrated to maintain a higher cash with an index of 0.13 against 0.08 of constrained firms.

In turn, with respect to the debt profile, we found that constrained firms had the most concentrated debt in the short term than the unconstrained firms. The first group had 61% of the debt in the short term and 39% long term, while the second got 47% in the short term and 53% long term.

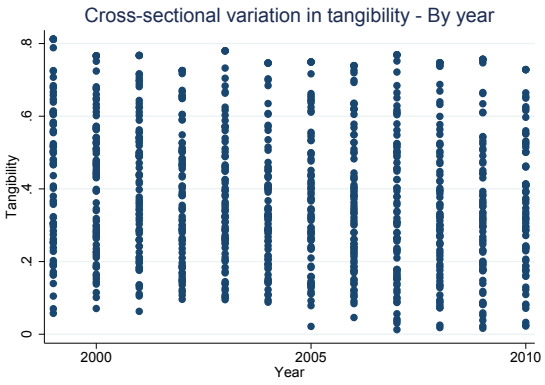
3.5.2 Firms' heterogeneity in tangibility

Figure 17 shows the cross-sectional variation of our baseline tangibility measure for the whole sample of firms, by year and just before the implementation of the "New Bankruptcy Law" in Brazil, that triggered the expansion cycle on credit and real estate prices. Figure 3a shows that the variation on firms' tangibility is observed across the years. Moreover, Figure 3b reveals that there is both a substantial degree of cross-sectoral variation on the average firm tangibility (SIC 2 Digits on manufacturing firms) and of cross-sectional variation on firms belonging to a given industry (similar pattern are found in East European countries, as shown in CAMPELLO; LARRAIN, 2016).

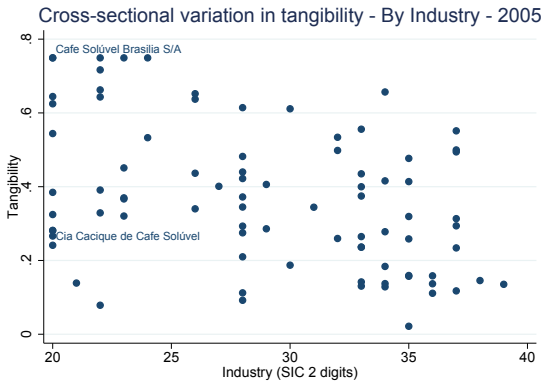
This observed cross-sectional variation in firms' tangibility within industries is crucial to our empirical strategy. By focusing on the high and bottom fraction of tangibility in firm-industry groups, our approach requires enough cross-sectional variation in the tangibility. By doing so, our goal is to control for common industry shocks that affect firms in a given industry, but that can

Figure 17: Tangibility (Fixed assets / total assets) distribution, by year and by industry (sic 2 digits)

(a) Tangibility variability - By year



(b) Tangibility variability - By industry



Source: Authors' elaboration.

be very different from industry to industry. For example, Brazilian food industry (SIC2=20) have benefited from both increasing local private consumption and increasing foreign demand for these goods, such as poultry meat, industrialized soy products, among others. However, other industries did not have the same positive macroeconomic shocks: Tobacco products (SIC2=21), for example, was prejudiced by the imposition of increasing restrictions on tobacco consuming. In other words, by focusing on firms with different tangibility in the same economic activity (industry), we control for specific industry shocks.

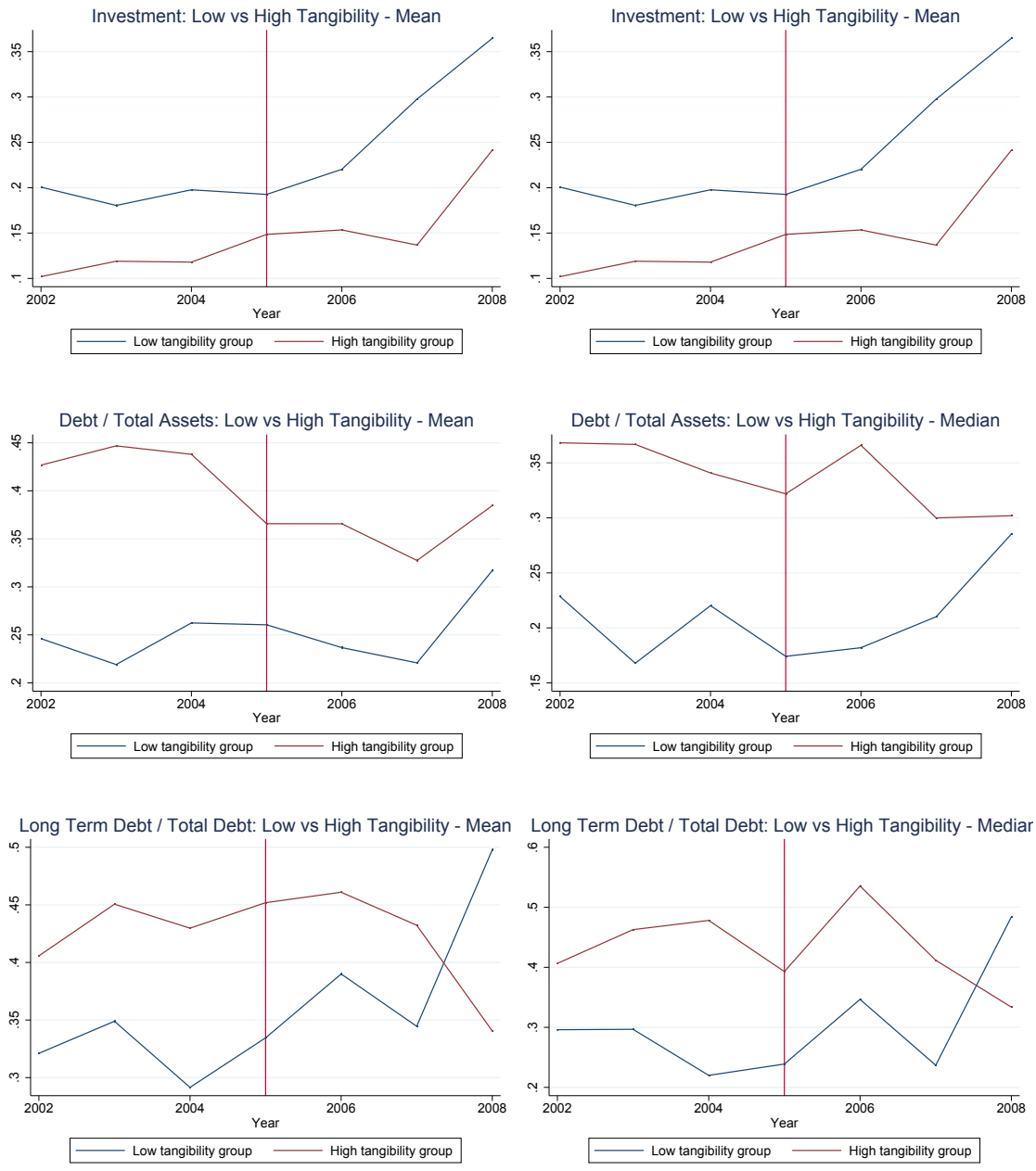
We highlight in Figure 17(b) two firms that are very similar in its activities and mix of products: Café Solúvel Brasília S/A and Cia Cacique de Café Solúvel. Both are coffee producers, which is an important item in the Brazilian food industry. We can plausibly argue that both firms are exposed to the same macroeconomic and investment opportunities shocks. However, one firm had in 2005 a fixed assets / total assets very different from the other. We explore these differences in the potential use of collateral to access external finance across companies to estimate the effect of the booming collateral prices on debt financing and investment.

3.5.3 Are there differences in investment and firm financing between the two groups?

Exploiting the observed cross-section variability in tangibility, we analyze mean and median of our three independent variables (investment, debt / total assets, and long-term debt / total debt), by period (before and after the expansion credit cycle) and by group (low vs high-tangibility).

In Figure 18, we show the evolution of these variables between the high and low-tangibility groups.

Figure 18: Mean and median of our core independent variables, by year and by low-high tangibility groups



Source: Authors' elaboration.

The graphical evidence suggests that the low-tangibility group increased more than proportionally its investment, total debt and long-term debt / total debt after 2005. If we look at the numbers (Table 19), we can see that the post-pre difference in investment is 5.3 p.p. and 8.3 p.p.

for the high and low-tangibility firms, respectively. The t-test for differences of means confirms that these differences are statistically significant at 1%. However, when we look at the difference of the differences (row and column "Difference" of Panels A, B and C in Table 19), we conclude that the control group (low-tang) increased its investment 3.0 p.p. more than the treatment group (high-tang). This difference is economically relevant (around 17.7% considering the pre-event investment mean of the low-tang group) albeit not statistically different of zero by the standard t-test. We find more prominent results in total debt (Panel B), where the control group increased in 6.6 p.p. its total debt -total assets ratio relative to the treatment group. This difference is statistically significant (5%) and economic sizeable: a 27% increase in leverage considering the pre-treatment level of the control group.

Table 19: Results of the difference-in-difference estimations, by variable and period, 2002-2008

Note: This table reports the results of the basic empirical strategy. We divide our financial constraints sample of firms into two groups, based on its within sector tangibility. Treatment group is formed by companies belonging to the top 40% of its industry's fixed assets / total assets. Control group consists of firms at the bottom 40% of its industry's fixed assets / total assets. Time variables are defined by the effectiveness of the Law 11,101/2005, the "New Bankruptcy Law" (it started to vigorate in Brazil by May 12th, 2005). Therefore, After refers to the period 2005 to 2008 and Before refers to the period 2002 to 2004. In all Panels, we calculate the average of each referred variable pre and post-law, i.e., we have the averages Before and After the event. Each Panel refers to a different dependent variable, as described above. The difference in differences in each Panel is the variable of interest. ***, **, and * implies significance at the 99%, 95% and 90% levels, respectively.

Variables	**Before**	**After**	**Difference**	**N**
	Panel A: Firm Investment			
Treat	0.102	0.154	0.053^{***}	326
Control	0.177	0.260	0.083^{***}	420
Difference			-0.030	
	Panel B: Total Debt / Total Assets			
Treat	0.400	0.355	-0.045^{*}	391
Control	0.244	0.265	0.021	482
Difference			-0.066^{**}	
	Panel C: Long Term Debt / Total Debt			
Treat	0.432	0.405	-0.027	386
Control	0.347	0.436	0.089^{***}	452
Difference			-0.116^{***}	

Source: Authors' elaboration.

One could also note that we find the larger difference between the two groups in the composition of debt. Before 2005, the average firm in the high-tangibility group had 43.2% of its debt in long-term debt (LTD). Meanwhile, the average low-tangibility firm had only 34.7% of its debt in LTD. After the bankruptcy law and the expansion cycle in the credit market, the former firms seem to have benefited the most of this new macroeconomic environment: a meaningful increase of 11.6 p.p. (33% of the pre-event average) relative to the high-tangibility group. This empirical evidence is particularly important because it implies one heterogeneous effect of the credit cycle on Brazilian firms: less tangible firms, potentially facing more financial frictions before the credit boom, effectively increased its long-term debt, which interest rates are normally lower and the higher maturity allows the firm to best manage its investment decisions over the business cycle.

3.5.4 Discussion of results

Our study starts from the division of the firms into two groups (financial constrained and financial unconstrained) according asset size criterion. From these groups we distinguish the first results, which include: (i) companies constrained had investment level lower than the unconstrained, validating anticipated by the literature (KIRCH; PROCIANOY; TERRA, 2014; BERNINI; GUILLOU; TREIBICH, 2016); (ii) the two groups showed similar degree of tangibility; (iii) the financial constrained firms had concentrated debt in the short term, while financial unconstrained firms had most of the long-term debt; (iv) the financial constrained firms presented a debt evolution much higher that the unconstrained firms. Regarding this last item, we understood that constrained firm, which had very little access to credit in the initial period, have access to credit easier from the credit expansion. Although this issue draws attention, it corroborates with results of others studies. In evaluating Mexico's industrial sector from 1984 to 1994, the authors found that, over a period of credit expansion, financial constraints appear to have been eased especially for smaller firms (GELOS; WERNER, 2002). The same results were found by Bernini, Guillou and Treibich (2016), through an investigation in the French financial market.

Later, in our sample of financially constrained firms (bottom 40% of Total Assets, in each year of our sample) we find that the low-tangibility group (bottom 40% of its industry fixed assets / Total assets) increased investment around 3.0 percentage points (pp) more than the high-tangibility group (top 40% of its industry fixed assets / Total assets), comparing means before (2002-2004) and

after (2005-2008) the credit expansion cycle. This variation is economic meaningful: 17% Increase in investment for the low-tangibility group relative to the high-tangibility group, considering the pre-event mean (17.7%).

Although some evidence is shown that the credit boom facilitated investment for firms with a low fraction of pledgeable assets, our results are stronger (statistically and economically) for firm financing. The average firm in the low-tangibility group increased total debt / total assets around 6.6 p.p. more (27% of increase considering the pre-event average) than the average firm belonging to the high-tangibility group, and this difference is also statistically significant (5%). This result suggests that collateral played a major role in firms access to debt before the credit expansion, and that the higher supply of credit that followed the New Bankruptcy Law allowed less tangible firms to increase access credit markets.

Besides the increase in total debt, an important finding of our empirical estimations is that the composition of debt changed significantly more to the low-tangibility group. The average low-tangibility firm increased the share of long-term debt on total debt after 2005 around 11.6 p.p. (33%) more relative to an average high-tangibility firm. More than statistically significant (1%), this result is economically sizable: a 1/3 increase in the fraction of long-term debt on total debt to low-tangibility firms relative to high-tangibility firms after 2005 indicates a qualitatively meaningful improvement on the access of corporate credit in Brazil, especially for firms that were potentially out of the market before the credit expansion cycle.

In summary, our study points to an easier access to credit for low-tangibility firms, which enabled the increasing the investment of this group's firms. This result ends up supporting the evidence of Guermazi (2014), who evaluated the credit expansion in Tunisia over the period of financial liberalization (1999 to 2005). The author found that, although assets' tangibility plays an important role in firms' investment behavior, the sensitivity of investment in relation to tangibility lost strength over the credit boom. On the other hand, such a conclusion is not definitive, and studies such as Gelos and Werner (2002) found that the importance of collateral in the credit decision did not reduce over the credit expansion observed in Mexico.

3.6 Conclusion

Our macro evidence shows that following the enactment of the New Bankruptcy Law (Law n. 11.101/2005), real estate prices and overall credit in Brazil indeed started to rise faster, especially to individuals. We exploit this new expansion cycle in both credit and real estate prices to estimate how firms' financing and investment reacted to this new economic environment. We sort firms accordingly to its ex-ante likelihood on being financially constrained, i.e., the extent to which a given firm faces difficulties to raise external finance. We than compare how firm financing and investment behaves relative to a counterfactual firm in the same industry (SIC 2 digits), in order to control for industry specific economic shocks.

In our sample of financially constrained firms (bottom 70% of Total Assets, in each year of our sample) we find that the low-tangibility group (bottom 40% of its industry fixed assets / total assets) increased investment around 3.0 percentage points (p.p.) more than the high-tangibility group (top 40% of its industry fixed assets / total assets), comparing means before (2002-2004) and after (2005-2008) the credit expansion cycle. This variation is economic meaningful: a 17% increase in investment for the low-tangibility group relative to the high-tangibility group, considering the pre-event mean (17.7%).

Although some evidence is shown that the credit boom facilitated investment for firms with a low fraction of pledgeable assets, our results are stronger (statistically and economically) for firm financing. The average firm in the low-tangibility group increased total debt / total assets around 6.6 p.p. more (27% of increase considering the pre-event average) than the average firm belonging to the high-tangibility group, and this difference is also statistically significant (5%). This result suggests that collateral played a major role on firms access to debt before the credit expansion, and that the higher supply of credit that followed the New Bankruptcy Law allowed less tangible firms to increase access credit markets.

Besides the increase in total debt, an important finding of our empirical estimations is that the composition of debt changed significantly more to the low-tangibility group. The average low-tangibility firm increased the share of long term debt on total debt after 2005 around 11.6 p.p. (33%) more relative to an average high-tangibility firm. More than statistically significant (1%), this result is economic sizable: an 1/3 increase in the fraction of long term debt on total debt to low-tangibility

firms relative to high-tangibility firms after 2005 indicates a qualitatively meaningful improvement on the access of corporate credit in Brazil, especially for firms that were potentially out of the market before the credit expansion cycle.

Taken together, our results suggest that the credit boom of the second half of the 2000s alleviated financial constraints in Brazil, especially for smaller, less tangible publicly traded firms, ending up to increase long term financing and boosting corporate investment.

Although we find consistent evidence for the multiplier effect lead by the collateral channel and exposed in Almeida and Campello (2007), our results suggest a primary role of banks alleviating collateral requirements, potentially because of the higher supply of credit and better growth opportunities in the economy. In this sense, our evidence supports the idea that the relaxing collateral requirements dominated the multiplier effect during the 2000s credit expansion in Brazil.

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