ESSAYS ON FISCAL POLICY AND PUBLIC DEBT

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TESE DE DOUTORADO
EM ECONOMIA

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Universidade de Brasília
Faculdade de Administração, Contabilidade e Economia
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TESE DE DOUTORADO SUBMETIDA AO PROGRAMA DE PÓS-GRADUAÇÃO EM ECONOMIA DA UNIVERSIDADE DE BRASÍLIA COMO PARTE DOS REQUISITOS NECESSÁRIOS PARA A OBTENÇÃO DO GRAU DE DOUTOR DOUTOR.

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Brasília/DF, Fevereiro de 2023.

FICHA CATALOGRÁFICA

CARVALHO CUNHA, DANIEL
Essays on Fiscal Policy and Public Debt. [Brasília/DF] 2023.
xxx, nnnp., 210 x 297 mm (ENE/FACE/UnB, Doutor, Tese de Doutorado, 2023).
Universidade de Brasília, FACE, Departamento de Economia.
Departamento de Economia
1. Regional Fiscal Multiplier   2. Cash transfer
3. Public Debt   4. Inflation Linked-Bonds
5. Green bonds   6. ESG
I. FACE/UnB   II. Fiscal Policy and Public Debt

REFERÊNCIA BIBLIOGRÁFICA

CARVALHO CUNHA, DANIEL (2023). Essays on Fiscal Policy and Public Debt. Tese de Doutorado, Publicação FACE Fevereiro/2023, Departamento de Economia, Universidade de Brasília, Brasília, DF, Fevereiro 2023.

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À Joana Russo Cunha, Chica e Fiona.

ACKNOWLEDGEMENTS

Agradeço à Professora Marina Delmondes de Carvalho Rossi, ao Professor Tomás Rodriguez Martinez e ao José Luiz Rossi Junior pelo suporte e pelas contribuições dados no decorrer da pesquisa. Aos meus colegas do FMI e da Secretaria do Tesouro Nacional, por darem o apoio necessário para que eu pudesse conciliar os estudos com a atividade profissional, em especial à Fabian Bornhorst, Joana Pereira, Frederik Toscani, Otavio Ladeira de Medeiros, Luiz Fernando Alves e Paulo Moreira Marques. Ao meu grande amigo Rodrigo Medeiros de Lima, pelo apoio, inspiração e debates ao longo dos anos que influenciaram a minha forma de pensar. Ao meu amigo Renan Abrantes de Sousa, pela parceria e ajuda ao longo do mestrado e doutorado. À minha família, pela compreensão pelos momentos que tive de dedicar aos estudos em vez de à família. Por fim, agradeço especialmente a Samia Marques Russo, minha esposa, que me apoiou e incentivou incondicional e incansavelmente.

ABSTRACT

On the fiscal policy, we estimate the subnational employment and GDP multiplier of Brazil's 2020 federal cash transfers to vulnerable households. Using two-stage least squares regressions we estimate a formal employment multiplier and then apply an analytical transformation to recover an implied GDP multiplier in the range of 0.5-1.5. The lower bound of this range lies below most estimates in the literature, which may result from the exceptional constraints imposed by the pandemic on supply chains and consumption. Nevertheless, even using the lower end of our range implies that federal cash transfers played an important role in supporting employment and GDP. On public debt, we study role play by sovereign Inflation Linked Bonds (ILBs) and Environmental, Sustainable, and Governance (ESG) bonds. About ILBs, we formally show that government bonds' term premia measures how much an external observer cannot learn about fundamentals from prices and the demand for public bonds will be higher if agents expect that the term premia will contract over timer and/or the variation of the term premia is low. Importantly, we analytically demonstrate that the demand for fixed rate bonds is positively impacted by the demand/ information of inflation linked bonds. Using a difference-in-differences approach, we empirically estimate the impact of the creation of a sovereign Inflation-Linked Bond (ILB) market finding that the opening leads to a significant improvement across different term premia metrics for EMs, but it is not significant for AEs. About sovereign ESG bonds, we explore a granular data base from the IDB covering 625 corporate and sovereign ESG bond issuances in the Latin America the Caribbean region (LAC) outstanding in offshore markets to investigate how a sovereign ESG bond issuance can boost the corporate ESG bond market. Using a difference-in-differences approach, we empirically estimate the impact of sovereign issuers tapping into the external ESG debt market finding that it roughly leads to a 50 percent increase in the volume of corporate bond issuances, and 25 percent increase in the number of ESG corporate bond issuances in the external market after two years. On the mechanisms, we argue that building a sovereign ESG market provides a benchmark enhancing the price discovery process of corporate bond issuances.

Keywords: Local fiscal multiplier, cash transfers multiplier, inflation linked-bonds, sovereign green and thematic bonds

RESUMO

Sobre a política fiscal, estimamos o multiplicador do PIB e do emprego subnacional das transferências federais de renda do Brasil em 2020 para famílias vulneráveis. Usando regressões de mínimos quadrados em dois estágios, estimamos um multiplicador de emprego formal e, em seguida, aplicamos uma transformação analítica para recuperar um multiplicador de PIB na faixa de 0,5-1,5. O limite inferior deste intervalo encontra-se abaixo da maioria das estimativas da literatura, o que pode resultar dos constrangimentos excecionais impostos pela pandemia às cadeias de abastecimento e consumo. No entanto, mesmo usando o limite inferior de nossa faixa, isso implica que as transferências de renda desempenharam um papel importante no apoio ao emprego e ao PIB. Sobre a dívida pública, estudamos o papel desempenhado pelos títulos soberanos indexados à inflação (ILBs) e pelos títulos ambientais, sustentáveis e de governança (ESG). Sobre ILBs, mostramos formalmente que os prêmios de prazo dos títulos do governo medem o quanto um observador externo não pode aprender sobre os fundamentos dos preços e a demanda por títulos públicos será maior se os agentes esperarem que os prêmios de prazo se contraiam ao longo do tempo e/ou da variação de os prêmios de prazo são baixos. É importante ressaltar que demonstramos analiticamente que a demanda por títulos prefixados é positivamente impactada pela demanda/informação de títulos indexados à inflação. Usando uma abordagem de diferença em diferenças, estimamos empiricamente o impacto da criação de um mercado soberano de títulos vinculados à inflação (ILB), descobrindo que a abertura leva a uma melhoria significativa em diferentes métricas de prêmios de prazo para econonomias emergentes. Sobre títulos ESG soberanos, exploramos uma base de dados granular do BID cobrindo 625 emissões de títulos ESG corporativos e soberanos na região da América Latina e Caribe (LAC) pendentes em mercados offshore para investigar como uma emissão de títulos ESG soberanos pode impulsionar o ESG corporativo mercado de títulos. Usando uma abordagem de diferenças em diferenças, estimamos empiricamente o impacto de emissores soberanos explorando o mercado externo de dívida ESG, descobrindo que isso leva a um aumento de aproximadamente 50 porcento no volume de emissões de títulos corporativos e um aumento de 25 porcento no número de Emissões de títulos corporativos ESG no mercado externo após dois anos. Sobre os mecanismos, argumentamos que a construção de um mercado ESG soberano fornece uma referência que aprimora o processo de descoberta de preço das emissões de títulos corporativos.

Palavras-chave: Multiplicador fiscal local, Multiplicador de transferências, títulos vinculados à inflação, títulos soberanos verdes e temáticos

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CHAPTER 1

ESTIMATING THE EMPLOYMENT AND GDP MULTIPLIERS OF EMERGENCY CASH TRANSFERS IN BRAZIL

1.1 INTRODUCTION

In response to the COVID-19 pandemic, the Brazilian federal government provided substantial fiscal and financial support to the economy. Policies included tax relief for companies and households, temporary income support to vulnerable individuals, credit facilities for firms (mostly small and medium sized companies, which as a group employ the largest share of the population), a job-retention scheme with subsides for furlough workers, and debt relief and cash transfers to states and municipalities, to name a few. In all, fiscal measures in 2020 had a direct impact on the primary deficit of over 7 percent of GDP. In addition, public banks opened pandemic-related credit lines worth close to 5 percent of GDP.

The Emergency Aid (henceforth EA) cash transfers program, which cost nearly BRL 300 billion (4.3 percent of GDP) in 2020, was the most emblematic pillar of the government's response package. Offering basic income to informal workers (employed or unemployed) and vulnerable households, the program was instrumental to stave off a dramatic rise in poverty and inequality which COVID-19 would have brought by its direct impact on labor markets, as shown by Flamini et al. (2021) and Cardoso (2020). For several months the EA more than offset labor income losses for households at the bottom four deciles of the income distribution, effectively increasing their real household income relative to pre-pandemic levels.

While the primary goal of the EA program was to provide social assistance in the exceptional context of the pandemic and forced social distancing, data on retail activity, mobility, and growth performance - as well as anecdotal evidence - suggests that the EA also provided an important cushion to the economy overall, avoiding a deeper recession in Brazil (with a contraction of around 4 percent of GDP, Brazil's recession was milder than roughly 75 percent of advanced economies and emerging markets).

In this paper we aim to formalize the impact of EA on growth. We use monthly private formal employment data at the municipal level to estimate the employment effects of the EA, taking advantage of the large heterogeneity in the distribution of the EA eligibility across Brazil's several thousand municipalities. As in Chodorow-Reich (2019), we first estimate an employment multiplier (in our case a formal employment multiplier - the number of private formal sector jobs created by each BRL 100 thousand disbursed by the federal government on EA) and then propose an analytical transformation to translate the employment multiplier into an output multiplier.

At the municipal level, no high-frequency data for the large informal sector (nearly 50 percent of the labor market) or public employment (5 percent) exists. This necessitates the choice to estimate a private formal sector employment multiplier rather than a total employment multiplier.[1] To obtain causal estimates, we instrument for EA spending with the pre-pandemic share of *Bolsa Familia* (longstanding

[1]Informality rates vary from 16 to 97 percent across municipalities.

conditional cash transfer program) recipients at the municipal level.

Chodorow-Reich (2020) shows that if geographic units are equal to the size of US states ($N = 1/50$) or smaller then factor mobility and the disturbance arising from non-treated regions (SUTVA-micro) should have a negligible impact on regional estimates. The size of Brazilian municipalities ($N = 1/5483$) provides an ideal practical setting for the use of regional data in macroeconomics as a shock to an minuscule area has no significant effect on any other single area. Furthermore, (formal) labor mobility is limited in Brazil. Dix-Carneiro and Kovak (2019) show evidence of imperfect interregional labor mobility after a negative labor demand shock (brought about by a trade policy reform). Dix-Carneiro and Kovak (2017) also find minimal effects of regional shocks on inter-regional worker mobility and Cavalcanti et al. (2019) find an important spatial segmentation of labor markets.

In the baseline estimations we find a formal employment multiplier around 0.5 (for the six-month window April-September 2020). This implies an annual cost per private formal job of over BRL 400,000 or USD 78,000. An important consideration, however, is the role of informality. In municipalities with low informality, estimating a private formal sector employment multiplier is similar to estimating a total employment multiplier. But in municipalities with high informality, the private formal sector employment multiplier is likely to be substantially below the total employment multiplier (in the limit, with informality at 100 percent, the private formal employment multiplier becomes meaningless and is equal to 0 as cost per job goes to ∞). Analyzing an expansion of the *Bolsa Familia* in 2009, Gerard et al. (2021) find that municipalities with a high informality rate have an underlying cost per formal job ten times larger than regions with low informality, meaning that the formal employment multiplier of informal municipalities is 1/10 of the ones with lower informality rate.

To account for this heterogeneity in the formal sector employment multiplier, we also run regressions interacting the EA variable with the pre-pandemic structural informality rate. Taking a weighted average of the resulting multipliers to obtain a national estimate yields a private formal employment multiplier over 6 months of around 1.6 - that is, a cost per year-job of BRL 124,000 (or USD 24,000). This national multiplier is larger than the one obtained in the baseline regression - as well as a multiplier obtained through a simple conditional mean (around 0.9) - because municipalities with a higher share of formal employment in total formal employment (i.e., with higher weight) tend to have larger multipliers.

We derive analytical expressions which allow us to transform the estimated formal employment multiplier into a total employment multiplier and a GDP multiplier, similar to Chodorow-Reich (2019) but taking into account the share of formal and informal workers and their relative productivity. This yields a range of 1-3.5 for the annual total employment multiplier (annual cost per job of USD 6,000-24,000) and a broad range of 0.3-1.8 for the GDP multiplier, with a preferred range of 0.5-1.5 considering most adequate specifications. Essentially, lower estimates are obtained from specifications without the informality interaction term, while the upper estimates originate from specifications with such interaction term.

Conceptually, our estimates yield a cross-region transfer multiplier. Pennings (2021) shows that a transfer multiplier should be smaller than a purchase multiplier, depending on how large the marginal propensity to consume (MPC) is and on how much of the transfers are being spent locally (how open the economy - here municipality - is). The higher the MPC and the less open municipalities are, the higher will be the transfer multiplier.[2]. In addition, how permanent transfers are also plays an important role, unless the share of hand-to-mouth households is very large. For the US, Pennings (2021) finds a cross-regional transfers multiplier of 1.5 for permanent transfers and 1/3 for one-off transfers.

The duration of the EA transfers was contingent on the persistence of the acute phase of the pandemic. Even considering that uncertainty on the unfolding of the outbreak was high, the expected length of time of the EA was neither a pure one-off nor a permanent transfer. We interpret, thought, that EA was closely related to temporary transfers, implying that our estimated impact works as a lower bound to a permanent cash transfer multiplier. The EA spanned from 2020 (4.3 percent of GDP) to 2021 (1 percent of GDP), but we focus on 2020 due to the provision of a better natural experiment of the discretionary fiscal policy impulse as the size of the EA was an unexpected shock, and the first six months of the program were somewhat isolated of the lagged effects of the monetary policy.

There exists only limited empirical evidence on the cross-regional transfer multiplier in emerging markets. Given the wide range of theoretically plausible multipliers, we see as one of the main contributions of our paper that it provides a benchmark for the case of Brazil. It is also among the first studies of the impact of Covid-related response programs on economic activity in emerging markets. Last, we build on the literature by carefully considering the role of informality in obtaining multipliers in an emerging market. Closely related studies include Sadoulet et al. (2001), who find a multiplier range of 1.6-2.5 analyzing the PROCAMPO program in Mexico, and Egger et al. (2019), who estimate a local cash transfer

[2]Considering a corner example in which the households of a given municipality have a "home-bias" preference to fully spend their transfers in locally-produced goods/services, then the transfer multiplier would be 1, similar to a situation in which the government directly purchases goods/services from this municipality.

multiplier of 2.6 using novel experimental evidence from Kenya. Using panel regressions Denes et al. (2018) estimate that the *Bolsa Familia* multiplier was as high as 4 in Brazil in the 2004-2010 period.[3]

Corbi et al. (2019) estimate formal employment purchase multipliers at the municipal level in Brazil. The authors exploit a discontinuity in the allocation of federal transfers to municipalities, based on population size thresholds, to study the impact of externally financed municipal government spending on formal employment over the period 1999-2014. They find a cost-per-annual-formal-job of USD 8,000-13,000 and a preferred GDP multiplier when accounting for the informal sector of 2.4.[4] This is at the very top of the range of direct spending (purchase) multipliers generally obtained in cross-sectional studies for the US. Chodorow-Reich (2019) cite a multiplier range of 1-2.5, with a preferred point estimate of 1.8, and Serrato and Wingender (2016) find a local multiplier between 1.7 and 2 using a Census shock to map expenditure changes that depend on the local population size.

In all, the empirical evidence for emerging markets has so far shown large local multipliers, with both purchase and transfer multipliers significantly above 1. The upper end of our estimated GDP multiplier range (0.5-1.5) corresponds to estimates previously found in the literature while the lower end lies substantially below most cross-sectional GDP multiplier estimates. Of course, the exceptional nature of the Covid-19 pandemic might plausibly explain lower estimates. It seems intuitive that forced social distancing and substantial restrictions on the supply of

[3] Estimating national transfer multipliers in emerging markets, Bracco et al. (2021) find a general multiplier of 0.9 for Latin America, compared to 0.3 for developed economies. The difference is mainly explained by the larger share of hand-to-mouth households in EMs economies.Neri et al. (2013) find an implied GDP multiplier of *Bolsa Familia* of 1.8.

[4] Corbi et al. (2019) calibrate the productivity ratio of informal to formal workers to $\rho = 0.55$. Updating their estimation using recent work by Ulyssea (2018), we obtain $\rho = 0.81$ and, thus, an implied GDP multiplier of 3.5.

various services, with a (likely associated) sharp increase in aggregate household savings, would lead to lower multipliers ceteris paribus[5].

We nonetheless interpret our estimated local EA multiplier as providing a lower bound for what the corresponding national multiplier would be.[6] Chodorow-Reich (2019) for instance argues that local spending multipliers lay out a rough lower bound for the aggregated national (deficit financed) spending multiplier for a closed economy when monetary policy is passive. Similar factors are at play in this study. Brazil is a relatively closed economy, and surely more closed than its individual municipalities. In addition, monetary policy was accommodative before the pandemic and relatively passive during the first six months of the EA.

When real rates do not rise in response to higher government spending, the standard multiplier measured at the national level rises, but this effect is "netted"out in cross-sectional regressions, as all municipalities are equally affected by monetary policy.[7] Bellifemine et al. (2022) highlight that non-tradeable employment and wealth are the key confounders of region monetary policy heterogeneity, and our estimated multiplier controls for these variables.

The remainder of this paper proceeds as follows. Section 1.2 provides more details on the EA program, motivating the subsequent discussion of data and research design in sections 1.3 and 1.4. Section 1.5 presents the results and section 6 presents brief concluding considerations. Additional tables and figures not presented in the

[5]Auerbach et al. (2020) provide evidence for the US that points in a similar direction. Looking at the average impact of Covid-related fiscal response measures in a large group of countries during 2020, Deb et al. (2021) also find that demand policies (including, though not restricted to, cash transfers) had less impact on economic activity during stringent lockdowns

[6]A national EA multiplier cannot be obtained given the very short time period under consideration.

[7]Chodorow-Reich (2019) considers the case when monetary policy response is restricted by a zero lower bound scenario instead. Similarly, real rates fall in that scenario following the increase in government spending.

main text can be found in the Appendix.

1.2 BACKGROUND: THE BRAZILIAN EMERGENCY AID CASH TRANSFER PROGRAM

The EA was a means tested program of monthly disbursements, which covered roughly 60 percent of the total workforce in the initial months. It was initially also very generous, providing an estimated replacement rate of 40 percent of the average income in the informal sector and increasing the real household income of the bottom four deciles of the income distribution by 20 percent (at least in May and June 2020) according to Flamini et al. (2021). The original design (Law no. 13.983, April 2, 2020) offered support during 2020Q2. Given the continued outbreak of COVID-19, however, the authorities extended the program twice in 2020, first through end 2020Q3 and later for the whole year (at that stage with tighter eligibility and a 50 percent reduction in stipends).

Eligibility. The EA initially offered a monthly basic income of BRL600 (and twice as much for single parents) to all contributors to Brazil's public social security system (INSS), participants of the national single registry (*Cadastro Único*), citizens registered as individual micro-entrepreneurs (MEI), and informal workers not registered in other federal assistance programs. In addition, *Bolsa Família*[8] beneficiaries could temporarily migrate to the EA program. The eligibility age was 18 years or older. Participants must belong to a household with per capita monthly income of no more than BRL 522 (half the minimum wage) or total monthly income up to BRL 3,135 (thrice the minimum wage). Finally, participants could not

[8] *Bolsa Família* was Brazil's most important social assistance program. Prior to the pandemic it covered around 14 million households, paying a monthly benefit of less than BRL 200, at a total annual cost of 0.4 percent of GDP.

have had annual taxable gross income greater than BRL 28.5 thousand in 2018.

Financial inclusion. A vast number of EA participants did not have a banking account at the time of the first transfer. Caixa Econômica Federal (CEF) thus offered digital banking accounts, debit cards, and cell phone apps to include them in the system. Participants with accounts in other institutions could decide which bank to use to withdraw the cash. Importantly, banks could not withhold EA transfers to citizens with outstanding debt or past due overdraft accounts. The social safety net got wider and digital.

Fiscal costs. The EA transferred approximately BRL 40 billion (1/2 percent of GDP) per month to recipients during 2020Q2 and 2020Q3. Owing to tighter controls over claimants' eligibility and reduction in monthly stipend, the cost of EA transfers fell to half the amount in 2020Q4.

1.3 DATA

Brazil is a federal republic with three levels of government - federal, state and municipal. The unit of analysis in this paper are Brazil's municipalities. Our main (independent) variable of interest is the amount of Emergency Aid (EA) transfers disbursed by the federal government (Ministry of Citizenship) to the public bank CEF, which in turn was responsible for disbursing payments to individuals.

Ideally, we would like to use a measure of total municipal employment (or even a municipal GDP proxy) as the dependent variable. However, no suitable GDP proxy is available, and total employment data (covering both the formal and informal sector) is not available at the municipal level, even at annual frequency.[9] The

[9] The continuous household survey PNAD provides data at the national and state level for

dependent variable is thus private formal employment as measured by the government's administrative CAGED dataset. To account for extreme outliers linked to likely measurement problems, we exclude the top and bottom 0.5 percent of municipalities in terms of employment growth and EA per capita.[10] Below we briefly summarize the main data sources. Table 2.2 presents summary statistics while Figure 1.2 shows correlations between the main variables.

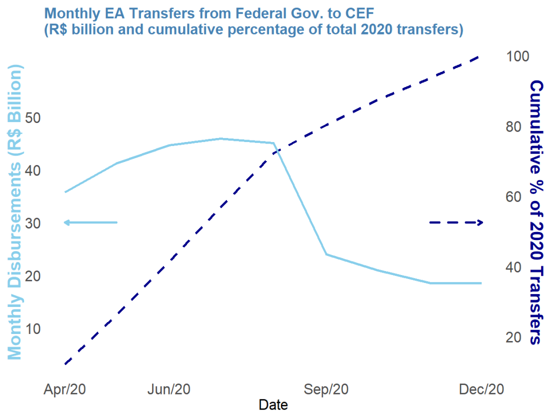
Emergency Aid: Data on EA payments at the municipal level are provided by the Brazilian Ministry of Citizenship. The first round of EA payments - six installments from April to September/2020 at the full original amount - represented 80 percent of total year-transfers (Figure1.1). For each round, new beneficiaries were informed *ex ante* about the payment schedule of future installments.

both the formal and informal sectors, but is not representative at the municipal level. Only the population census - conducted once every ten years - provides a full picture of employment at the municipal level.

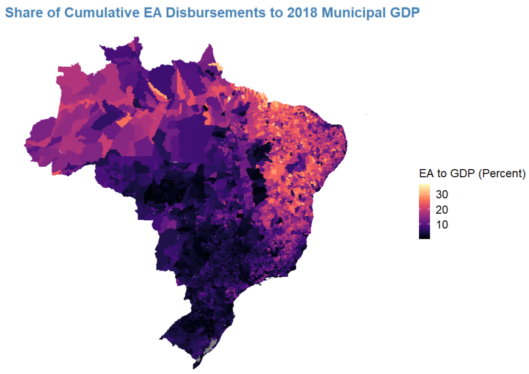
[10] Regression results are not meaningfully different when outliers are retained, but parameter stability decreases somewhat.

Figure 1.1: EA Disbursements Over Time and Municipal Heterogeneity

(a) Monthly Disbursements



(b) Regional (municipal) Distribution



Formal Employment: CAGED is an administrative database that covers in principle the universe of private formal workers in Brazil. There was a significant change in its methodology at the end of 2019 and, thus, we chiefly focus on the series from 2020 onward to ensure consistency. We use CAGED's November 2021 vintage which offers the most up to date payroll series, after incorporating ex-post notifications of hiring/layoffs and correcting methodological issues.

Mobility: The google mobility index tracks more than 2,400 Brazilian Municipalities. We use the average of mobility indices for groceries and pharmacies, parks, retails and recreation, transit stations, and workplaces.

COVID-19 Deaths: New daily Covid-19 related deaths per million were computed using the administrative data from the Brazilian Ministry of Health.

2010 Census-based Indicators: Informality at the municipal level is defined as the ratio of informal employment to total employment (considering the occupied population over 10 years of age). The share of non-white population covers all individuals that did not declare themselves as white or Asian. Urbanization is calculated as the ratio of urban population to total population in each municipality. Agriculture as a share of total employment captures employment in agriculture as a share of total employment. Last, services as a share of total employment were computed taking into account the sectors outside of the perimeter of agriculture, manufacturing, and public administration.

Bolsa Familia Recipients: The share of the population receiving *Bolsa Familia* benefits in each municipality was calculated using administrative data from the Ministry of Citizenship, as of December 2019.

Job Protection Program: The Emergency Benefit for Preserving Employment and Income (BEm) was launched in April 2020, allowing for temporary reductions in working hours or contract suspension in the formal sector, by mutual agreement between employers and employees. The program backed workers by partially compensating for the associated salary losses, in an amount proportional to the unemployment insurance to which the employee would have been entitled to if she lost the job (i.e., pro-rated by the percentage reduction in working hours). Importantly, each employee could have more than one BEm agreement, either because she worked for more than one firm or for agreeing first to a cut in working hours and later to a contract suspension, or vice versa. The BEm data, from the Ministry of Economy, shows the number of monthly agreements aggregated at the municipal level.

Pandemic-related credit to SMEs: Besides the EA and BEm, pandemic-related lending facilities were the third largest pillar of the government's fiscal response. Among those, the Emergency Program for Access to Credit (PEAC), launched in June 2020, was the largest pandemic-related credit line to the real sector ($\approx$ 1.2 of the GDP). The program unlocked credit to SMEs, associations, private foundations, and cooperatives with underlying 2019 revenues between BRL360,000 and BRL300 million. The federal government provided guarantees to the credit lines, covering 80 percent of the face value of each operation. The PEAC data, from the Brazilian Development Bank (BNDES), shows monthly disbursements aggregated at the municipal level.

Table 1.1: Summary Statistics

Statistic	N	Mean	St. Dev.	Min	Median	Max
Δ Formal Employment per Thousand	5,483	−0.4	13	−203.6	−0.3	348
Emergency cash transfers (EA) per capita	5,483	1,120	322	212	1,118	2,016
Share of *Bolsa Familia* Receivers	5,483	9	6	0	7	27
Share of non-white Population	5,472	52	24	1	57	99
Bolsa Familia (BF) Cover Ratio	5,472	1.08	0.23	0.22	1.09	2.90
Job Agreements per Formal Worker	5,333	0.31	0.28	0	0.23	9
PEAC Disbursements per capita	3,843	154	314	0.0	44	6,309
Δ Covid-19 Deaths per million	5,483	2	2	0	2	18
Urbanization Rate	5,479	64	22	4	65	100
Agriculture as Share of Employment	5,479	35	18	0	35	87
Services as Share of Employment	5,479	53	12	8	53	87
Informality Rate	5,478	64	17	16	67	97
GDP per capita (Thousand)	5,483	23	24	5	17	575
Population (Thousand)	5,483	38	225	1	12	12,325

Note: The unit of analysis is municipalities, the lowest administrative level in Brazil. Data on formal employment creation, EA cash transfers, and number of job protection agreements under the BEm program refer to the sum over April-September 2020. The share of the population receiving *Bolsa Familia* benefits in each municipality was calculated using administrative data from the Ministry of Citizenship, as of December 2019. The *Bolsa Familia* cover ratio is defined/estimated as the number of poor households receiving *Bolsa Familia* benefits as of end-2012 as a share of the estimated number of poor households according to the 2010 Census. Covid-19 deaths are measured as the average of daily new deaths over April -September 2020. GDP per capita is taken from the 2018 municipal accounts. At the average 2020 exchange rate to the USD, mean GDP per capita of BRL 23,000 corresponds to about USD 4,000. Population figures are collected from the Brazilian Institute for Geography and Economics (IBGE). We exclude the top and bottom 0.5 percent of municipalities in terms of the change in formal employment and emergency aid per capita.

1.4 RESEARCH DESIGN

1.4.1 Baseline Econometric Specification and Identification

To estimate the cross-sectional *formal* employment multiplier we follow a standard approach, as set out for example in Chodorow-Reich (2019). Specifically, we regress the total change in private formal employment per capita between April and September 2020 on total EA disbursements per capita at the municipal level over the same period - which is the period with the bulk of EA disbursements.[11]

[11] In robustness exercises we change the window of analysis to the initial disbursement period (April-June) or the full 9-month period of EA disbursements (April-December).

By design, EA disbursements per capita are determined by the share of the local population which is eligible for the program. Given that eligibility at the individual (and, thus, municipal) level was determined based on data as of end-2019, it could be viewed as largely exogenous. However, selection into or out of the program can give rise to endogeneity concerns. In particular, households might either opt to not seek the EA even when eligible and non-eligible households could find a way around the proposed targeting, with the latter having occurred especially at the beginning of the program according to reports. We thus instrument for EA disbursements with the share of the population receiving conditional cash transfers under the *Bolsa Familia* program *pre-pandemic*. As explained above, being a *Bolsa Familia* recipient is one of the criteria for being eligible for EA disbursements. Figure 1.2 presents the correlations between the main variables used in the analysis. As a first takeaway, we note that there is a high correlation between EA disbursements per capita and the share of pre-pandemic *Bolsa Familia* recipients at the municipal level, already suggesting that the inclusion restriction for the instrument is likely to be satisfied.

A key difference between *Bolsa Familia* and EA relies on the fact that budgtary constraints were binding for *Bolsa Familia* but not for EA. Gerard et al. (2021) indicate that the total number of *Bolsa Familia* beneficiaries is set nationally with specific rules guiding the allocation of recipients across municipalities. Conversely, the EA, because of the exceptionally status granted by the state of calamity issued in 2020, did not have a de facto national cap neither a competition for resources within the EA budget envelope across municipalities. As a result, at the municipality level the *Bolsa Familia* has more exogenous features than the EA. Accordingly, Gerard et al. (2021) provide further evidence on exogenous aspects of

Bolsa Familia finding that the 2009 expansion of the program had causal impact on formal employment growth.

There is a body of literature - e.g. Barrientos et al. (2016) and Ribas et al (2011)- pointing out the exogenous features of *Bolsa Familia* figures at the municipal level. According to Barrientos et al. (2016), even though the selection of *Bolsa Familia* receivers depends on their per capita household income, the placement at the municipal level depends mostly on the pre-programme poverty level of the municipality. As such, the pre-pandemic *Bolsa Familia* program assignment can be considered exogenous concerning the EA transfers and control variables.

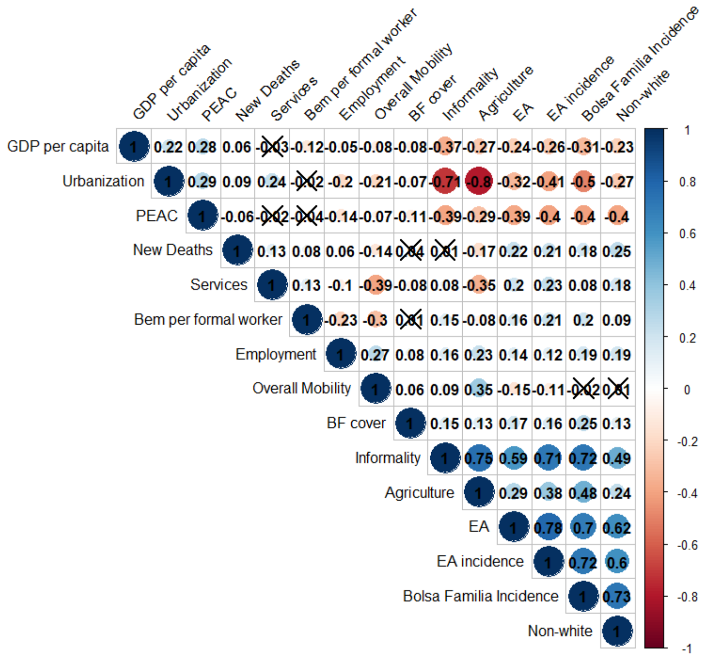
We estimate the following set of equations:

$$\Delta e\hat{a_{h,m}} = \alpha_S + \phi_0 + \phi_1 BF_m + \phi_2' X_m + \xi_m \tag{1.1}$$

$$\Delta fe_{h,m} = \alpha_s + \mu + \beta \Delta e\hat{a_{h,m}} + \gamma' X_m + \varepsilon_m \tag{1.2}$$

where α_s and α_S are state fixed effects, $\Delta f\hat{e_{h,m}} = \sum_{t=0}^{h} e\hat{a_{h,m}}$ denotes the monthly change in formal employment per capita from period 0 to h for a municipality m, and $\Delta e\hat{a_t^m} = \sum_{t=0}^{h} e\hat{a_t^m}$ indicates the sum of EA transfers per capita from period 0 to h for a municipality m. BF_m stands for the share of individuals entitled to *Bolsa Familia* benefits pre-pandemic, X_m represents the set of control variables which in the baseline specification comprises the daily new COVID-19 deaths per million and formal employment trends (by taking into account the change in formal employment in Q1 2020 and in 2019). ε_m and ξ_m are robust standard errors. Last, we control for the municipal informality rate, given the direct link between formal job creation and the structural level of informality.

Figure 1.2: Correlation Matrix



Note: The matrix shows correlations for the variables summarized in Table 1. Correlations marked with "x" are not significant at 95 percent confidence levels.

1.5 RESULTS

1.5.1 The (Formal) Employment Multiplier

Table 1.2 presents the results from the baseline specification (column (1)) as well as a number of robustness exercises and extensions. Specifically, column (2) adds the share of services in total employment as a control variable, column (3) controls for the urbanization rate, and column (4) includes google mobility as a control variable. The first stage F-statistic is highly significant across specifications (Table 1.10 in the Appendix shows the first stage results).

In the baseline specification we find a formal employment multiplier of 0.53 -

in other words 100,000 of EA payments create 0.53 private formal jobs for a 6 month period. This implies an annual cost-per-formal-job of BRL 378,813 =(100,000/0.528)*2. At the average 2020 USD-BRL exchange rate, this corresponds to a cost per job of around USD 73,000. Comparing this with the simple OLS results presented in Table 1.9 of the Appendix, we see that the 2SLS multiplier estimate is around three times as large as the OLS one. This is consistent with the bias discussed in section 1.4.1, whereby worse pandemic outcomes and thus worse labor market outcomes would lead to higher EA per capita.

It is instructive to map the estimated coefficient for the formal employment multiplier to a total employment multiplier. This will allow to gain a fuller picture, facilitate comparisons with the literature, and obtain a GDP multiplier estimate, which we'll discuss in section 1.5.2.

The private (formal) employment multiplier β_{FE} is defined in the standard way as the change in private formal employment d_{FE} for a given change in government spending d_G. The total employment multiplier is defined equivalently and the expression can be rearranged as follows

$$d_E = \beta_E d_G \implies \beta_E = \frac{\frac{d_E}{E_t}}{\frac{d_G}{E_t}} \tag{1.3}$$

where E_t denotes the total number of jobs. Multiplying and dividing by $\frac{d_{FE}}{FE_t}$ and collecting terms gives

$$\beta_E = \Theta \beta_{FE} \frac{1}{\omega_{fe}} \tag{1.4}$$

where Θ is the elasticity of total employment to private formal employment and ω_{fe} is the formality rate. The ratio $\frac{1}{\omega_{fe}}$ aims to adjust the private formal employment multiplier by its relative size to total employment, considering the informal and public (formal) employment shares. We recover $\omega_{fe} = 0.45$ from the 2010 census data, which has the most comprehensive account of employment formality (using the number from the household survey PNAD[12], 0.46, would not make a big difference, however), yielding an adjustment factor of $\frac{1}{\omega_{fe}}$=2.22.

The elasticity of total employment to private formal employment, Θ, can be expressed as the sum of three underlying components:

$$\Theta = \omega_{fe} + \Psi\omega_{if} + \eta\omega_{pe} \tag{1.5}$$

.

$\omega_{if} = 0.5$ is the informality share and $\omega_{pe} = 0.05$ is the share of public employment. Additionally, we take into account the elasticity of informal employment to private formal employment, denoted by $\Psi = 2.35$[13] and obtained from PNAD (which includes monthly data on both formal and informal employment) for the same 6-month window used in the regressions (April-September 2020). We assume that the elasticity of public employment to changes in private formal employment, given by η, is zero considering that the public sector dynamics were most likely driven by the healthcare response and, thus, orthogonal to EA transfers. As a result, we find that Θ equals to 1.63.

Multiplying the estimated β_{FE} as set out in equation (1.4) yields a total employment multiplier of 1.6 for the 6-month window. In other words, around 1.6 jobs

[12]We do not rely on PNAD data for our regressions given that the PNAD survey is not representative at the municipal level).

[13]The elasticity was computed using realized seasonal adjusted series from PNAD. From 2019 to 2020 (after the labor reform of 2017/2018 was implemented) the mean of Ψ has been relatively stable around 2.4, with an underlying coefficient of variation close to 0.8

(0.5 formal and 1.1 informal) were generated for each BRL 100,000 of paid EA. Equivalently, the annual cost-per-job is around BRL 103,000, or USD 20,000.

Sectoral employment structure as a possible confounder While the baseline specification controls for state fixed effects and important structural municipal characteristics such as informality, one additional concern might be that a different local economic structure (for example in terms of sectoral composition) is correlated with the share of pre-pandemic *Bolsa Familia* recipients, while also impacting the sensitivity of formal employment to the pandemic shock. One plausible mechanism might be that structurally poorer municipalities have a larger services sector which in turn suffered more during the pandemic. At first sight the correlations shown in Figure 1.2 do not suggest that this is a particularly pronounced correlation, but in a robustness exercise we nevertheless include the share of services employment as a control[14]. Importantly, the employment share in the service sector shields our analysis from potential concerns due to the heterogeneous impact of monetary policy across regions as Bellifemine et al. (2022) point that non-tradeable employment is a key confounder of the heterogeneity. In addition, we control for the urbanization rate; another potentially important structural municipal characteristic which could impact how the pandemic affected employment creation[15].

Looking across columns (2) and (3) of Table 1.2, the estimated formal employment multiplier drops marginally when adding service employment as a control variable and drops somewhat more to 0.44 when also adding the urbanization rate, but

[14]We also ran exercises controlling for employment in industry and in commodity-sensitive sectors, which yielded similar results to the ones shown in Table 1.2.

[15]Bellifemine et al. (2022) argue that wealth is a relevent confounder of monetary policy heterogeneity across regions. Bloom et al. (2008) show that urbanization rate and wealth are highly correlated. As such, we proxy the stock of wealth by the urbanization rate given the lack of stock of wealth data in Brazil.

remains significant. Both additional control variables have the expected sign, with more services-intensive and urbanized municipalities experiencing weaker job growth.

Controlling for Google Mobility Google mobility data - for which we would like to control given the pandemic-induced variability in mobility and, hence, economic activity - is only available for 2,210 (slightly less than half of all) municipalities. However, these 2,210 municipalities account for 89 percent of national GDP, 93 percent of total formal employment, and 77 percent of total EA disbursements. Including mobility in the regressions thus has two effects - (i) a composition effect, whereby we exclude municipalities which are on average smaller, poorer, more informal, more rural and more dependent on agriculture, and (ii) the direct impact of adding the control variable for a constant sample.

The change in the estimated multiplier is large when including Google mobility data, roughly doubling in the restricted sample (column (4)). As mentioned above, municipalities in the sample for which mobility data is available are larger and richer than those without mobility data. The change in the coefficient is entirely due to this composition effect and not because of the importance of mobility as a control variable. To formally test this, we re-run the regressions with the reduced sample but without mobility as a control, concluding that this only marginally changes the coefficient. The coefficient on mobility has the expected sign, with higher mobility associated with larger formal employment creation.

Intuitively, whether mobility data, as captured by Google, is available or not for a specific municipality suggests important differences in its level of development or other characteristics, beyond the control variables we include. These differences could be perhaps along dimensions we cannot observe such as internet connectivity, smart phone usage and related factors which could determine how much remote work, for example, is feasible. We investigated the difference in coefficients between the full and restricted sample further (including by controlling for additional observable variables which differ between the two groups) but could not obtain a clear explanation - beyond the hypothesis that there might be a true difference in treatment effect based on some unobservable sample characteristic. We continue to refer to the results for the full sample as the baseline, but given that the restricted sample accounts for around 90 percent of national GDP and formal employment we do not discard the higher estimates it yields.

Interacting pre-pandemic informality with EA As mentioned above, there exists a direct relationship between the informality rate in a municipality and the number of formal jobs created per capita over any time period - a marginal formal job in a highly informal municipality likely requires a larger change in local economic activity. This is an important concern for our research design, which might be even more important during the pandemic since informal and formal jobs were affected at varying degrees by lockdowns and social distancing. To allow for a difference in treatment effect between more and less formal municipalities we include an interaction term between the pre-pandemic informality rate and the EA disbursement per capita (instrumented by the interaction between the pre-pandemic share of *Bolsa Familia* recipients and the pre-pandemic informality rate).

Table 1.3 repeats the order of Table 1.2 in terms of the control variables included in each column but adds the interaction term between EA and pre-pandemic informality to all specifications. The national formal employment multiplier is obtained by taking a weighted average of the sum of the main effect and interaction

Table 1.2: Baseline Regression Results

	Cumulative Change in Formal Employment per capita			
	Instrument: Share of Population Receiving Bolsa Familia (Pre-pandemic)			
	Baseline	Adding Services Empl.	Adding Urbanization Rate	Adding Mobility
	(1)	(2)	(3)	(4)
EA per capita (BRL 100K)	0.528^{***}	0.443^{***}	0.415^{***}	0.666^{**}
	(0.138)	(0.138)	(0.135)	(0.267)
Covid-19 Deaths	0.002	0.063	0.123	0.261
	(0.115)	(0.115)	(0.112)	(0.185)
Informality Rate	-0.0005	-0.0001	-0.006^{**}	-0.010^{***}
	(0.002)	(0.002)	(0.003)	(0.004)
Δ Formal Employment Q12020	-0.436^{**}	-0.439^{**}	-0.436^{**}	-0.270
	(0.220)	(0.218)	(0.218)	(0.236)
Δ Formal Employment 2019	0.007^{***}	0.004^{*}	0.003	0.002
	(0.002)	(0.002)	(0.002)	(0.006)
Share of Services in Employment		-0.011^{***}	-0.008^{***}	0.002
		(0.002)	(0.002)	(0.003)
Urbanization Rate			-0.007^{***}	-0.010^{***}
			(0.001)	(0.002)
Overall Mobility				0.026^{***}
				(0.003)
Constant	-0.007^{***}	0.0003	0.006^{***}	0.005
	(0.002)	(0.002)	(0.002)	(0.003)
States Fixed effects	Yes	Yes	Yes	Yes
Implied Number of Year-Formal Jobs Created	633,557	531,614	498,444	799,007
Implied Cost per Year-Formal Job (BRL)	378,813	451,454	481,497	300,372
Implied Cost per Year-Formal Job (USD)	73,413	87,491	93,313	58,211
Implied Number of Year-Jobs (Formal and Informal) Created	2,287,845	1,919,720	1,799,939	2,885,306
Implied Cost per Year-Job (USD)	20,329	24,228	25,840	16,120
First Stage F statistic	417^{***}	406^{***}	400^{***}	91^{***}
Wu-Hausman	12.53^{***}	8.53^{***}	6^{**}	2
Observations	5,478	5,478	5,478	2,210
Residual Std. Error	0.013 (df = 5446)	0.013 (df = 5445)	0.013 (df = 5444)	0.011 (df = 2175)

*p<0.1; **p<0.05; ***p<0.01

Note: The unit of analysis is municipalities, the lowest administrative level in Brazil. Data on formal employment creation, and EA transfers refer to the sum over April-September 2020. Google mobility and Covid-19 deaths are measured as the average over April-September 2020. The share of population receiving *Bolsa Familia* is as of December/2019.

effect across all municipalities (using the municipalities' share in national formal employment as the weights).

The coefficient on the interaction term is highly significant and negative, indicating that more informal municipalities create less formal jobs for each BRL 100,000 of EA payment - an intuitive result. More interestingly, the average weighted formal employment multipliers we obtain are significantly higher than the ones in Table 1.2. The fact that more formal municipalities with a higher multiplier also account for a large share of total formal employment, leads to a national cross-sectional multiplier which is larger than the simple average relationship implied. Figure 1.3 makes this point visually by showing that municipalities with a higher multiplier also have a larger weight. Columns (1)-(4) show private formal employment multipliers of 1.5-1.7, implying an annual cost per formal job of BRL 132,000-117,000 (USD 23,000-26,000). The implied total employment multipliers are in the range of 5-6, with an annual cost per job as low as BRL 30,000 (around USD 6,000).

Instrumenting EA transfers using municipal ethnic patterns and BF cover ratio As an additional robustness exercise, we use the share of non-white population (NWP) measured by the 2010 Census and the estimated *Bolsa Familia* cover ratio of poor households to instrument EA transfers per capita[16]. Historically, the NWP has limited access to higher education in Brazil[17] and thus larger unemployment rates than white Brazilians. Figure 1.2 shows a high correlation between EA disbursements per capita and the share of non-white population at the municipal signalling the potential strength of the instrument. Aiming to capture some idiosyncratic components of municipal poverty features that are fully

[16]We also considered using either NWP or the BF cover ratio as single instrument to the EA transfers. However, considering the interaction with informality, the combined use of NWP and BF cover outperformed the isolated specifications as a feasible instrument.

[17]Mello (2021) provides a detailed discussion on this topic.

mapped by ethnics patterns, we use as an additional instrument the estimated *Bolsa Familia* cover ratio of poor households. The ratio takes into account the number of *Bolsa Familia* beneficiaries at end-2012 and the estimated number of poor Households according to 2010 Census[18].

Table 1.4 is built such as Table 1.2 with the difference that EA transfers per capita are instrumented by the share of NWP instead of the share of *Bolsa Familia* receivers. The formal employment multiplier is, on average, twice as large as that detailed in Table 1.2. Moreover, the results shown by Table 1.4 are in line with Table 1.3, pointing to a multiplier around 1.2. The tests reported in Table 1.4 provide evidence of the strength of our instruments, indicating absence of endogeneity or over-identification issues.

Interacting pre-pandemic informality with EA using alternative instruments As highlighted before, it is useful to estimate the formal employment multiplier directly taking into account the formality level of different municipalities. Thus, like in Table 1.3, we add to Table 1.4's specifications an interaction term between the pre-pandemic informality rate and the EA disbursements per capita (instrumented by the interaction between the pre-pandemic share of NWP and the pre-pandemic informality rate, keeping the instruments already used in Table 1.4[19]).

Similar to Table 1.3, Table 1.5 shows the national formal employment multiplier

[18]Rougier et al (2018) argue that the municipal-level cover ratio of *Bolsa Familia* is a good indicator of the municipality's capacity to identify, enroll and register poor individuals. As a result, the authors claim that the ratio is a good exogenous predictor of the cross-municipality variation of *Bolsa Familia* to GDP ratio.

[19]We ran regressions instrumenting the interaction of informality with EA by the interaction of the BF cover ratio with the informality rate. The implied formal employment multiplier is about 25 percent larger but in all specifications we reject the null hypothesis of the Sargan Test that the model is not over-identified.

taking a weighted average of the sum of the main effect and interaction effect across all municipalities. On average, the obtained formal employment multiplier is 20 percent larger than in Table 1.3.

Controlling for the government's Job Protection Program (BEm) and pandemic-related credit to SMEs (PEAC) As discussed in section 1.3, BEm was explicitly designed to limit formal sector job losses by allowing for a flexible reduction in working hours with partial income compensation. Trying to explicitly identify the impact of BEm on formal employment runs into an endogeneity problem that is significantly worse still than that of assessing the impact of the EA - municipalities in which the pandemic had a large impact would have seen higher selection into BEm participation. At the same time, the direct and mechanical impact of the BEm has to be an improvement in formal job dynamics relative to a counterfactual without BEm. PEAC credit lines face a similar endogeneity issue as the selection into the program was likely impacted by local economic conditions.

Given this difficulty, we do not aim to retrieve the impact of BEm and PEAC on employment (even though it is a very interesting question in its own right) and therefore did not include them as regressors in our baseline specifications. Nevertheless, given the direct link between BEm and formal employment and the impact of pandemic-related credit on formal employment, we investigate how our results change when BEm and PEAC are indeed controlled for. We opted for a specification in which BEm is normalized by total formal employment - in essence, the share of jobs protected by the program.

Table 1.6 adds the share of formal jobs covered by BEm as a control variable: column (1) adds the BEm variable to a specification otherwise identical to column (1) of Table 1.2, while columns (2), (3), and(4) do the same for column (1) of

Tables 1.3, 1.4, and 1.5 respectively.

The estimated formal employment multiplier drops significantly relative to the baseline without BEm and PEAC, by over 40 percent in the specification without the informality interaction (column (1)), by around 25 percent in the version with the informality interaction (column (2)) and when instrumenting EA transfers by ethnic patterns (column(3)), and about 10 percent in the specification instrumenting EA transfers by ethnic patterns and the BF cover ratio (column (4)). The specifications with BEm and PEAC as a control variables thus provide us with the lower end of our estimated multiplier range. Note that the (non-causal) coefficient of the BEm variable is highly significant and negative, suggesting that the negative selection effect dominates the positive mechanical association with formal employment retention.

Figure 1.3: Formal Employment Multiplier and Formal Employment Weight by Municipalities

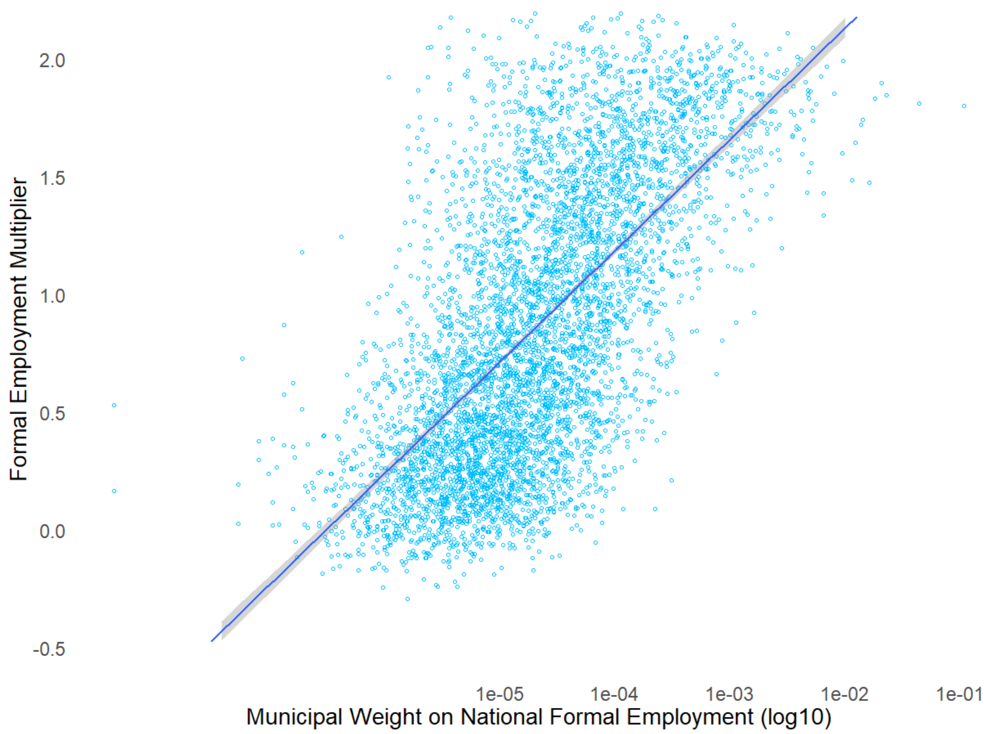


Table 1.3: Interaction with Informality Regression Results

	Cumulative Change in Formal Employment per capita			
	Instruments: Pre-pandemic Share of Population Receiving Bolsa Familia (BF) and BF*informality			
	Baseline	Adding Services Empl.	Adding Urbanization Rate	Adding Mobility
	(1)	(2)	(3)	(4)
EA per capita (BRL 100K)	3.118***	3.096***	3.003***	3.380***
	(0.434)	(0.433)	(0.425)	(0.648)
EA per capita (BRL 100K)*Informality	−3.491***	−3.594***	−3.502***	−3.919***
	(0.520)	(0.522)	(0.517)	(0.843)
Covid-19 Deaths	−0.152	−0.087	−0.026	−0.018
	(0.118)	(0.118)	(0.115)	(0.195)
Informality Rate	0.033***	0.034***	0.028***	0.026***
	(0.005)	(0.005)	(0.005)	(0.008)
Δ Formal Employment Q12020	−0.434**	−0.432**	−0.257	
	(0.221)	(0.218)	(0.218)	(0.240)
Δ Formal Employment 2019	0.005**	0.001	0.0003	−0.003
	(0.002)	(0.002)	(0.002)	(0.006)
Share of Services in Employment		−0.012***	−0.010***	−0.003
		(0.002)	(0.002)	(0.003)
Urbanization Rate			−0.006***	−0.009***
			(0.001)	(0.002)
Overall Mobility				0.026***
				(0.003)
Constant	−0.031***	−0.024***	−0.017***	−0.016***
	(0.004)	(0.004)	(0.004)	(0.005)
States Fixed effects	*Yes*	*Yes*	*Yes*	*Yes*
Formal Employment Multiplier - Weighted Average Across Munis	1.624***	1.558***	1.504***	1.703***
	(0.497)	(0.499)	(0.487)	(0.643)
Implied Number of Year-Formal Jobs Created	1,949,041	1,870,596	1,805,242	2,044,448
Implied Cost per Year-Formal Job (BRL)	123,137	128,301	132,946	117,391
Implied Cost per Year-Formal (USD)	23,863	24,864	25,764	22,750
Implied Number of Year-Jobs (Formal and Informal) Created	7,038,204	6,754,930	6,518,928	7,382,730
Implied Cost per Year-Job (USD)	6,608	6,885	7,134	6,300
First Stage F statistic (EA)	1259***	1272***	1310***	201***
First Stage F statistic (EA*informality)	1531***	1568***	2122***	196***
Wu-Hausman	28***	23***	21***	18***
Observations	5,478	5,478	5,478	2,210
Residual Std. Error	0.013 (df = 5445)	0.013 (df = 5444)	0.013 (df = 5443)	0.011 (df = 2174)

*p<0.1; **p<0.05; ***p<0.01

Note: The unit of analysis is municipalities, the lowest administrative level in Brazil. Data on formal employment creation, and EA transfers refer to the sum over April-September 2020. Google mobility and Covid-19 deaths are measured as the average over April-September 2020. The share of population receiving the *Bolsa Familia* benefit is as of December/2019. To compute the national formal employment multiplier, we calculate individually for all municipalities the implied effect for their specific level of informality and then calculate the average by weighting with formal employment.

Table 1.4: Regression Results Using Ethnic Patterns and BF cover as Instruments

	Cumulative Change in Formal Employment per capita			
	Instruments: Non-white Share of the Population and BF Cover			
	Baseline	Adding Services Empl.	Adding Urbanization Rate	Adding Mobility
	(1)	(2)	(3)	(4)
EA per capita (BRL 100K)	1.312***	1.263***	1.217***	1.158**
	(0.302)	(0.300)	(0.296)	(0.525)
Covid-19 Deaths	−0.054	0.003	0.066	0.170
	(0.110)	(0.110)	(0.109)	(0.214)
Informality Rate	−0.005*	−0.005*	−0.011***	−0.014**
	(0.003)	(0.003)	(0.003)	(0.006)
Δ Formal Employment Q12020	−0.548***	−0.551***	−0.547***	−0.274
	(0.171)	(0.169)	(0.169)	(0.241)
pre_e_2019	0.006***	0.002	0.001	0.002
Δ Formal Employment 2019	0.006***	0.002	0.001	0.002
	(0.002)	(0.002)	(0.002)	(0.006)
Share of Services in Employment		−0.011***	−0.008***	0.001
		(0.002)	(0.002)	(0.003)
Urbanization Rate			−0.007***	−0.010***
			(0.001)	(0.002)
Overall Mobility				0.028***
				(0.004)
Constant	−0.012***	−0.005**	0.001	0.004
	(0.002)	(0.002)	(0.002)	(0.004)
States Fixed effects	Yes	Yes	Yes	Yes
Implied Number of Year-Formal Jobs Created	1,574,509	1,515,295	1,460,698	1,389,937
Implied Cost per Year-Formal Job (BRL)	152,428	158,385	164,305	172,669
Implied Cost per Year-Formal Job (USD)	29,540	30,694	31,842	33,463
Implied Number of Year-Jobs (Formal and Informal) Created	5,685,727	5,471,898	5,274,744	5,019,216
Implied Cost per Year-Job (USD)	8,180	8,500	8,817	9,266
First Stage F statistic	198***	198***	201***	68***
Wu-Hausman	18***	17***	15***	4 **
Sargan	2.86 *	1.32	3.3 *	1
Observations	5,472	5,472	5,472	2,208
Residual Std. Error	0.013 (df = 5440)	0.013 (df = 5439)	0.012 (df = 5438)	0.011 (df = 2173)

*p<0.1; **p<0.05; ***p<0.01

Note: The unit of analysis is municipalities, the lowest administrative level in Brazil. Data on formal employment creation, and EA transfers refer to the sum over April-September 2020. Google mobility and Covid-19 deaths are measured as the average over April-September 2020. The share of NWP comes from the 2010 Census. The estimated *Bolsa Familia* cover ratio of poor households takes into account the number of *Bolsa Familia* beneficiaries at end-2012 and the estimated number of poor households according to the 2010 Census.

Table 1.5: Interaction with Informality Regression Results (Alternative Instruments)

	Cumulative Change in Formal Employment per capita			
	Instruments: Non-white Share of Population(NWP), BF cover and NWP*Informality			
	Baseline	Adding Services Empl.	Adding Urbanization Rate	Adding Mobility
	(1)	(2)	(3)	(4)
EA per capita (BRL 100K)	4.073***	4.281***	4.056***	3.554***
	(0.679)	(0.680)	(0.672)	(0.926)
EA per capita (BRL 100K)*Informality	−4.350***	−4.769***	−4.480***	−4.071***
	(0.761)	(0.768)	(0.760)	(1.074)
Covid-19 Deaths	−0.188	−0.133	−0.070	−0.044
	(0.117)	(0.116)	(0.115)	(0.229)
Informality Rate	0.040***	0.045***	0.037***	0.027**
	(0.007)	(0.007)	(0.007)	(0.011)
Δ Formal Employment Q12020	−0.543***	−0.546***	−0.543***	−0.263
	(0.172)	(0.169)	(0.170)	(0.243)
Δ Formal Employment 2019	0.003	−0.001	−0.002	−0.003
	(0.002)	(0.002)	(0.002)	(0.006)
Share of Services in Employment		−0.013***	−0.010***	−0.004
		(0.002)	(0.002)	(0.004)
Urbanization Rate			−0.006***	−0.009***
			(0.001)	(0.002)
Overall Mobility				0.026***
				(0.004)
Constant	−0.040***	−0.034***	−0.027***	−0.017***
	(0.006)	(0.006)	(0.006)	(0.006)
States Fixed effects	*Yes*	*Yes*	*Yes*	*Yes*
Formal Employment Multiplier - Weighted Average Across Munis	2.212***	2.240***	2.139***	1.812***
Implied Number of Year-Formal Jobs Created	2,654,545	2,688,926	2,566,991	2,174,648
Implied Cost per Year-Formal Job (BRL)	90,410	89,254	93,494	110,362
Implied Cost per Year-Formal (USD)	17,521	17,297	18,119	21,388
Implied Number of Year-Jobs (Formal and Informal) Created	9,585,858	9,710,012	9,269,689	7,852,897
Implied Cost per Year-Job (USD)	4,852	4,790	5,017	5,922
First Stage F statistic (EA)	132***	132***	134***	47***
First Stage F statistic (EA*informality)	187***	185***	181***	52***
Wu-Hausman	19***	19***	17***	9***
Sargan	3.7*	1.8	3.6*	0.9
Observations	5,472	5,472	5,472	2,208
Residual Std. Error	0.013 (df = 5439)	0.013 (df = 5438)	0.012 (df = 5437)	0.011 (df = 2172)

*p<0.1; **p<0.05; ***p<0.01

Note: The unit of analysis is municipalities, the lowest administrative level in Brazil. Data on formal employment creation, and EA transfers refer to the sum over April-September 2020. Google mobility and Covid-19 deaths are measured as the average over April-September 2020. The share of population receiving the *Bolsa Familia* benefit is as of December/2019. To compute the national formal employment multiplier, we calculate individually for all municipalities the implied effect for their specific level of informality and then calculate the average by weighting with formal employment.

Table 1.6: Controlling for The Job Support Program (BEm) and Pandemic-related Lending (PEAC) Regression Results

	Cumulative Change in Formal Employment per capita			
	Instruments:			
	Share of Population Receiving BF	BF and BF*informality	Non-white Share of Population(NWP) and BF cover	NWP, BF cover, and NWP*Informality
	(1)	(2)	(3)	(4)
EA per capita (BRL 100K)	0.260	2.468***	1.265***	3.420***
	(0.198)	(0.484)	(0.403)	(0.851)
EA*Informality		−2.928***		−3.484***
		(0.597)		(0.912)
Number of BEm Agreements per formal worker	−0.010***	−0.009***	−0.009***	−0.009***
	(0.002)	(0.002)	(0.002)	(0.002)
PEAC per capita (BRL 100K)	0.132	0.203	0.193	0.249
	(0.176)	(0.176)	(0.172)	(0.175)
Covid-19 Deaths	0.122	−0.058	−0.022	−0.166
	(0.143)	(0.146)	(0.160)	(0.176)
Δ Formal Employment Q12020	−0.120	−0.113	−0.115	−0.112
	(0.183)	(0.184)	(0.187)	(0.187)
Δ Formal Employment 2019	0.007*	0.004	0.006	0.003
	(0.004)	(0.004)	(0.004)	(0.004)
Constant	−0.003	−0.023***	−0.009***	−0.030***
	(0.002)	(0.004)	(0.003)	(0.007)
States Fixed effects	*Yes*	*Yes*	*Yes*	*Yes*
Formal Employment Multiplier - Weighted Average Across Munis	0.260	1.218***	1.264***	1.930***
	(0.141)	(0.410)	(0.292)	(0.675)
Implied Number of Year-Formal Jobs Created	311,474	1,461,114	1,516,800	2,315,573
Implied Cost per Year-Formal Job (BRL)	770,529	164,258	158,227	103,646
Implied Cost per Year-Formal Job (USD)	149,327	31,832	30,664	20,086
Implied Number of Year-Jobs (Formal and Informal) Created	1,124,767	5,276,245	5,477,333	8,361,791
Implied Cost per Year-Job (USD)	41,352	8,815	8,491	5,562
First Stage F statistic (EA)	1125***	619***	186***	77***
First Stage F statistic (EA*Informality)		789***		125***
Wu-Hausman	2	22***	11***	10***
Sargan			2	2
Observations	3,789	3,789	3,785	3,785
Residual Std. Error	0.012 (df = 3755)	0.012 (df = 3754)	0.012 (df = 3751)	0.012 (df = 3750)

*p<0.1; **p<0.05; ***p<0.01

Note: The unit of analysis is municipalities, the lowest administrative level in Brazil. Data on formal employment creation, and EA transfers refer to the sum over April-September 2020. Google mobility and Covid-19 deaths are measured as the average over April-September 2020. The share of population receiving the *Bolsa Familia* benefit is as December/2019. Services as a share of total employment, urbanization rate, and the informality rate come from the 2010 census.

1.5.2 The Implied GDP Multiplier

Analytical Transformation

Our approach to obtain a GDP multiplier from the estimated (formal) employment multiplier closely follows Chodorow-Reich (2019), adjusting for the fact that we estimate a private formal employment multiplier rather than a total employment multiplier. The adjustment for the formal/informal dimension takes inspiration from the approach taken in Corbi et al. (2019).

Starting from the definition of formal employment multiplier, we derive (Y denotes the level of GDP and FE denotes the level of formal employment):

$$d_{FE} = \beta_{FE} d_G \implies \frac{d_{FE}}{FE_t} = \beta_{FE} \frac{Y_t}{FE_t} \frac{d_G}{Y_t} \tag{1.6}$$

The GDP multiplier β_Y, in turn, is equivalently defined as:

$$d_Y = \beta_Y d_G \tag{1.7}$$

which can be rearranged as

$$\beta_Y = \frac{\frac{d_Y}{Y_t}}{\frac{d_G}{Y_t}} \tag{1.8}$$

Now consider a production function $Y_t = A(N_t L_t)^{1-\alpha}$, where L denotes the stock of effective units of labor, and N indicates hours worked per worker. We define $L = FE + \rho IE + PE$ where IE equals informal employment, PE indicates public employment, and ρ is the relative productivity ratio of informal to formal workers. After totally differentiating the production function we obtain:

$$\frac{d_Y}{Y_t} = (\frac{d_N}{N_t} + \frac{d_L}{L_t})(1 - \alpha) \tag{1.9}$$

Dividing (1.9) by $\frac{d_{FE}}{FE_t}$ and rearranging terms yields

$$\frac{d_Y}{Y_t}\frac{FE_t}{d_{FE}} = (\chi + \Theta^L)(1 - \alpha) \tag{1.10}$$

where χ is the elasticity of hours per worker (N) to private formal employment (FE), and Θ^L is the elasticity of effective labor units (L) to FE, which is formally given by

$$\Theta^L = \omega^L_{fe} + \rho\Psi\omega^L_{if} + \eta\omega^L_{pe}. \tag{1.11}$$

Ψ and η are the elasticities of informal and public employment, respectively, to FE, while ω^L_{fe}, ω^L_{if} and ω^L_{pe} are the shares of private formal, informal, and public employment in effective labor.

Equation (1.11) is similar to equation (1.5), which defined the elasticity of total employment to formal employment Θ. However, since Θ^L maps the response of effective labor units, total employment is 'normalized' by the productivity ratio of informal to formal workers (ρ). Furthermore, we can rewrite the parameters ω^L_{fe}, ω^L_{if} and ω^L_{pe} as a function of their corresponding shares in total employment, as follows:

$$\omega^L_{fe} = \frac{E_t}{L_t}\omega_{fe} = \frac{1}{1 - (1 - \rho)\omega_{if}}\omega_{fe} \tag{1.12}$$

$$\omega^L_{if} = \frac{E_t}{L_t}\rho\omega_{if} = \frac{\rho}{1 - (1 - \rho)\omega_{if}}\omega_{if} \tag{1.13}$$

$$\omega_{pe}^{L} = \frac{E_t}{L_t}\omega_{pe} = \frac{1}{1-(1-\rho)\omega_{if}}\omega_{pe} \tag{1.14}$$

When ρ equals to 1, the underlying weight components of L_t are equivalent to their respective shares in total employment given that $E_t = L_t$, and, thus, $\Theta = \Theta^L$.

Plugging in equation (1.11) into (1.10) yields

$$\frac{d_Y}{Y_t}\frac{FE_t}{d_{FE}} = (\chi + \omega_{fe}^{L} + \rho\Psi\omega_{if}^{L} + \eta\omega_{pe}^{L})(1-\alpha) \tag{1.15}$$

Multiplying and dividing (1.8) by $\frac{d_{FE}}{FE_t}$ and finally combining (1.15), (1.6), and (1.8), yields:

$$\beta_Y = (1-\alpha)(\chi + \omega_{fe}^{L} + \rho\Psi\omega_{if}^{L} + \eta\omega_{pe}^{L})\frac{Y_t}{FE_t}\beta_{FE} \tag{1.16}$$

For a given private formal employment multiplier β_{FE}, and initial ratio of output per private formal worker $\frac{Y_t}{FE_t}$, the GDP multiplier β_Y increases with the labor share $(1-\alpha)$, with the elasticity of hours worked, informal employment or public employment to FE, and/or when private formal employment has a larger relative weight in effective labor L, either directly (higher share in total employment) or due to higher relative productivity (higher ρ).

It is worth noting the five differences with the transformation for the total employment multiplier obtained by Chodorow-Reich (2019). First, we have an explicit expression for the formality rate in the multiplicative factor. Second, we have an additional term in the multiplicative factor which captures the elasticity of informal employment to private formal employment adjusted for their relative productivity and the informality share. Third, we consider an extra term which measures the response of public employment to changes in private formal employment adjusted by the weight of public employment. Forth, we multiply by output per private formal worker rather than output per worker. Lastly, we use the elasticity of hours per worker to *formal* employment instead of *total* employment.

We demonstrate that the estimated GDP multiplier obtained from a total employment multiplier as a starting point is equivalent to the GDP multiplier derived from a private formal employment multiplier[20]. Furthermore, we show that when the informality rate is equal to zero we fall back to the exact equation defined in Chodorow-Reich (2019) given that ω_{if}^{L} goes to zero and $L_t = E_t$. Intuitively, the differences stem from the fact that we only observe a partial employment response, and the leap from the estimated private formal employment multiplier to a GDP multiplier is thus somewhat larger.

Calibration

As discussed in section 5.1, the continuous household survey PNAD allows us to observe the change in both informal and formal employment at the national level. We thus have a sense of how private formal, informal and public employment developed in both absolute and relative terms during the period of analysis - informal employment fell by more than formal employment, leading to a decrease in the informality rate, and public employment remained broadly stable[21].

We use the PNAD survey to calibrate Ψ, finding that for the 6 months period from April-September 2020, Ψ (seasonally adjusted) was equal to 2.35[22]. $\omega_{if}^{census} = 0.5$ is the weighted municipal informality rate from the 2010 Census. Based on the

[20] See Lemma 1 in the Analytical Appendix for a detailed discussion.

[21] Differently from the private sector, there are several legal protections granted to government employees implying that layoffs are not easily enacted. This imposes a downward rigidity to the level of public employees.

[22] The other parameters of the production function are calibrated using pre-pandemic data due to the lack of high frequency data to estimate unobservable variables.

work of Ulyssea (2018), we estimate ρ to 0.81.[23] In line with Corbi et al. (2019), we calibrated the labor share, $(1-\alpha)$, to 2/3 and the elasticity of hours per worker to total employment, χ^E, to 0.12. As a result, the elasticity of hours per worker to the stock of effective units of labor can be written as $\chi^L = \frac{L_t}{E_t}\chi^E = 0.11$, implying that the elasticity of hours per worker to the stock of employment is given by $\chi = \chi^L \Theta^L \approx 0.15$. As we mention before, we set $\eta \equiv 0$ considering the orthogonality of the reaction of public sector labor to EA transfers.

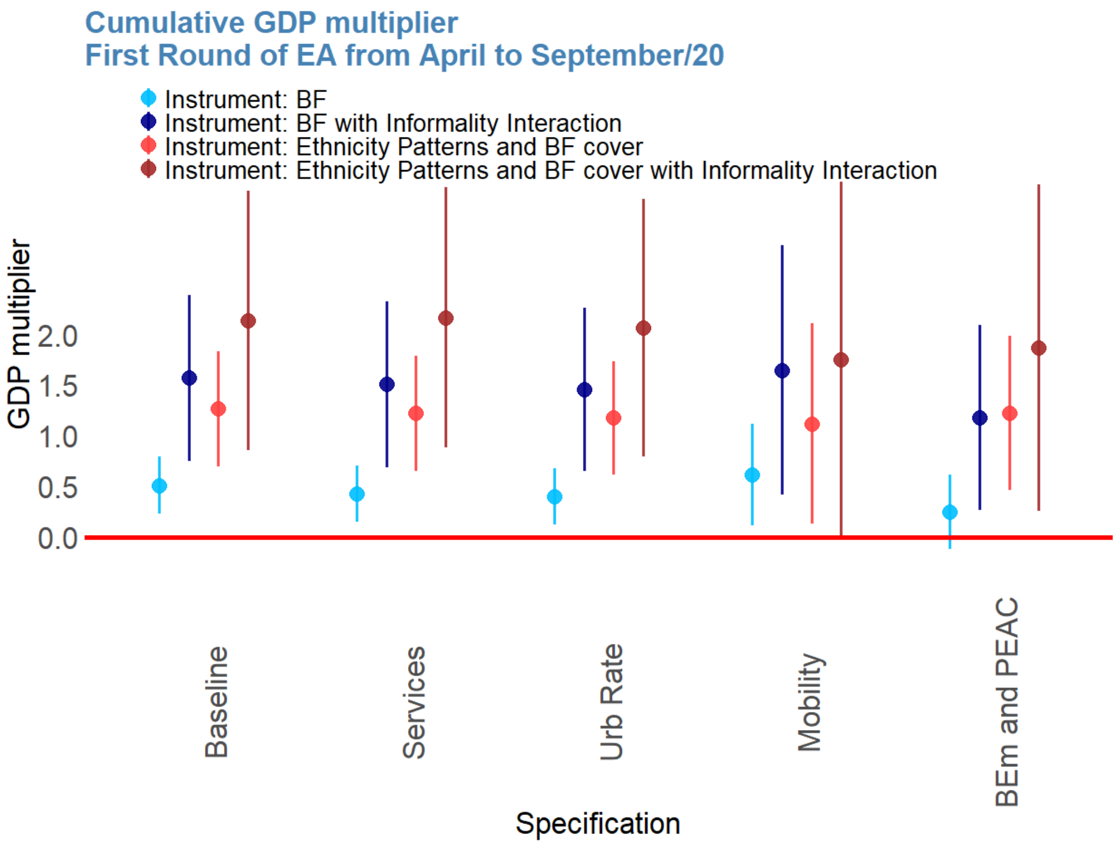
Results

Figure 1.4 shows the implied GDP multiplier retrieved from columns (1)-(4) of Tables 1.2, 1.3, 1.4, 1.5, and 1.6. The graph combines uncertainty both from the formal employment multiplier point estimate - which stems from the range obtained from different regression specifications - as well as the uncertainty from the confidence interval around each multiplier point estimate.

Incorporating the full range of uncertainty implied by the confidence intervals, we obtain a very wide range of multipliers,g from close to 0 to above 2. But focusing on the point estimates, a relatively consistent picture emerges. A GDP multiplier around 0.5 is obtained in specifications which use data from the full sample of municipalities and do not include the informality interaction. Instead, the estimated multiplier is around 1.5 when informality is accounted for by its interaction with the EA or when instrumenting EA transfers by ethnic patterns and the BF cover ratio.

[23] We proxy the productivity ratio of informal to formal workers (ρ) by the relative share of high skilled workers in the informal and formal sectors, using Ulyssea (2018)'s estimates for the Brazilian economy.

Figure 1.4: Implied GDP Multiplier



Note: $\beta_Y = (1-\alpha)(\chi + \omega^L_{fe} + \rho\Psi\omega^L_{if})\frac{Y_t}{FE_t}\beta_{FE}$ is the baseline transformation from formal employment to GDP multiplier. $(1-\alpha) = 2/3$ is the labor share, $\chi = 0.15$ is the elasticity of hours per worker to formal employment, $\Psi = 2.35$ is the elasticity of informal to formal employment, ω^L_{fe}=0.5 and $\omega^L_{if} = 0.45$ are municipal private formal and informality rates derived from Census 2010, $\rho = 0.81$ is the productivity ratio of informal to formal workers, Y_t is the 2020 GDP, and FE_t is the stock of formal workers. The vertical lines represent a 95 percent confidence interval.

Importantly, the above results are derived from regressions which estimate the relationship between the EA and formal employment creation for the six month window between April-September 2020. As discussed in section 1.3, this seems the most natural window for the analysis given that it covers the bulk of EA disbursements. Nevertheless, we re-run the baseline specification with informality interaction (column (1) in Table 3) for both a three-month window (April-June 2020) and a nine-month window (April-December 2020, covering the full EA disbursement in 2020). Table 1.7 presents the implied formal employment and GDP multipliers for these different time windows. The implied GDP multiplier (for the whole year) is broadly stable across estimation windows, but the falling employment multiplier suggests that the impact on economic activity faded over time.[24] Taking this result at face value, a possible explanation would be that transfers were incident on liquidity constrained consumers (and so had a large immediate effect) at the pandemic onset. However, as the economy recovered there were fewer liquidity constrained households and so the transfers had a smaller multiplier even though they became better targeted. The relaxing of liquidity constraints throughout the pandemic is in line with the implied saving rates from our multipliers.

Table 1.7: Formal Employment and GDP Multiplier at Different Horizons

	Cumulative change in employment per capita		
	EA time Horizon:		
	3 months	6 months	9 months
	(1)	(2)	(3)
Formal Employment Multiplier	2.655***	1.624***	1.117**
	(0.596)	(0.434)	(0.468)
GDP Multiplier	1.568***	1.569***	1.485**
	(0.352)	(0.419)	(0.622)
Implied Number of Year-Jobs (Formal and Informal) Created	7,380,900	6,820,800	6,500,940
Implied Cost per Year-Job (USD)	6,301	6,819	7,154
Baseline Controls	*Yes*	*Yes*	*Yes*
Ψ PNAD	3.27	2.35	2.01
First Stage F statistic (EA)	967***	1,259***	1,046***
First Stage F statistic (EA*Informality)	1,154***	1,531***	1,202***
Wu-Hausman	56***	28***	9***
Observations	5,474	5,478	5,478
Residual Std. Error	0.010 (df = 5441)	0.013 (df = 5445)	0.014 (df = 5445)

*p<0.1; **p<0.05; ***p<0.01

Note: The unit of analysis is municipalities, the lowest administrative level in Brazil. Data on formal employment creation, and EA transfers refer to the sum over April-June 2020 (3 months), April-September 2020 (6 months), and April-December 2020 (9 months). The elasticity of informal to formal employment, Ψ, for each time horizon is computed using seasonally adjusted numbers from PNAD.

[24] The falling employment multiplier shown in the first row and the stable GDP multiplier are consistent in the sense that the GDP multiplier throughout the paper refers to an annual concept while the employment multiplier refers to the estimation window, e.g. 3-months in column (1)).

Further discussion

As discussed in the introduction, from a theoretical perspective a large set of cross-sectional transfer multipliers is plausible ex ante. Here we use our empirical estimate and compare it to a theoretical multiplier obtained from parameter estimates taken from the literature. From a new Keynesian model standpoint for a closed economy, Pennings (2021) shows that when monetary policy is passive, the local transfer multiplier collapses to the following equation: $\beta_Y = \frac{\omega_T \alpha}{1-\omega\alpha}$ where ω is the share of hand-to-mouth households, ω_T is the fraction of transfers targeted at those hand-to-mouth households, and α is the degree of home bias (a measure of how closed the economy is).

The targeting of transfers and the share of hand-to-mouth households, while by no means fully observable, can be gauged to some degree from the literature. Bracco et al. (2021) estimate a share of hand-to-mouth households around half for Brazil ($\omega \approx 1/2$). Flamini et al. (2021) show that in the initial months of EA disbursements, the bottom half of the income distribution (which we can loosely assume here are the hand-to-mouth households) received around 75 percent of total disbursements, yielding $\omega_T \approx 3/4$. As a first step, we thus take ω and and ω_T as given and ask what α would be consistent with our multiplier estimate. Taking our estimated range of 0.5-1.5 for the multiplier, we obtain a range of 0.5-1 for α, suggesting relatively closed local economies. Previous evidence for Brazil does indeed point towards rather closed municipal economies given (i) the large share of non-tradable services in the Brazilian economy, especially in poorer municipalities, and (ii) limited (formal) labor mobility. Of course, the values for ω and ω_T might not be correctly measured and, moreover, parameters such as the share of hand-to-mouth households measured outside pandemic times might not accurately capture the dynamics during the pandemic.

One way to cross-check some of the intuition on how closed local economies are is to compare multipliers at the municipal level to those at a higher level of aggregation. Using state level data would lead to under-powered regressions. We thus exploit the fact IBGE provides an intermediate level of aggregation between municipalities and states, so called microregions. Given possible spillovers between neighboring municipalities, one would expect a higher multiplier for the microregion level regressions.

We compare the municipal level multiplier to the microregion level multiplier in two exercises. First, we focus on the restricted sample of municipalities for which we can observe (and, thus, control for) social mobility. Second, we consider the full sample of municipalities (not controlling for mobility). Since at the microregion level we have mobility data for all units of analysis, controlling for mobility does not give rise to the sample selection effect which arises at the municipal level.

Column (1) in Table 1.8 shows the GDP multiplier at the municipal level when controlling for mobility, while column (3) shows the municipal multiplier without controlling for mobility. Columns (2) and (4) show the corresponding microregion level multipliers. In line with intuition on the impact of openness, the estimated GDP multiplier increases somewhat at the microregion level. Furthermore, by comparing columns (2) and (4) one concludes that controlling for mobility per se has only a small impact on the estimated GDP multiplier. Therefore, to gauge the effect of using more aggregated regional data to estimate the GDP multiplier, the comparison between columns (3) and (4) - which avoids composition effects - seems most informative. The estimated multiplier increases from 1.45 in column (3) to 1.85 in column (4), a non-negligible difference which suggests that indeed some spillovers between neighboring municipalities occur.

Table 1.8: Additional Robustness tests: Changing the Unit of Analysis

	Controlling for Mobility		Not Controlling for Mobility	
	Municipalities	Microregion	Municipalities	Microregion
	(1)	(2)	(3)	(4)
GDP Multiplier	1.646***	1.733***	1.451***	1.852***
	(0.628)	(0.722)	(0.412)	(0.749)
Observations	2,208	547	5,478	547

*p<0.1; **p<0.05; ***p<0.01

Note: For columns (1) and (3), the unit of analysis is municipalities, the lowest administrative level in Brazil. For columns (2) and (4), the unit of analysis is micro regions. Data on formal employment creation, and EA transfers refer to the sum over April-September 2020. Columns (1) and (2) use the specification detailed in column (4) of Table 3. Columns (3) and (4) use the specification detailed in column (3) of Table 3.

All in all, our estimated multiplier range is consistent with a plausible set of underlying parameters, especially when considering forced reductions in consumption due to lock-downs reducing the marginal propensity to consume.

1.6 CONCLUSION

We provide an estimate of the GDP impact of Brazil's emblematic Emergency Aid (EA) cash transfer program, implemented from the outset of the Covid-19 pandemic. To the best of our knowledge, ours is among the first studies to focus on the output effects of fiscal response policies during this period. Although there is considerable uncertainty around the exact multiplier, our preferred specifications imply that it falls in the range of 0.5-1.5. This is somewhat lower than estimates found in the related literature for the pre-Covid period - both for the US and EMs, notably Brazil -, possibly reflecting the effect of lockdowns and social distancing on supply chains and consumption opportunities (forced savings). We also find that the impact of the EA was strongest in the first three months, when liquidity constraints were perhaps more pervasive. Still, even when using our most conservative estimates, the results suggest that the EA played an important role in cushioning the downturn and facilitating a rapid recovery. The counter-factual without EA would have been one with at least one million formal sector jobs and two million total jobs less, while 2020 GDP would have fallen by at least 2 percentage points more.

Looking ahead to further work, while progress had been made on understanding the size and heterogeneity of different types of multipliers in emerging markets, more analysis is needed to allow policy makers to design policies in the most growth friendly and inclusive way.

1.7 ANALYTICAL APPENDIX

Lemma 1. Equivalence of the GDP multiplier derived from private formal employment and total employment multiplier

Let β_E^L denote the total employment multiplier adjusted for effective units of labor, χ^L represent the elasticity of hours per worker to the total stock of effective units of labor, and α, L_t and Y_t represent the same parameters described by equation (1.9). Then, we can recover the GDP multiplier shown by equation (1.16) as follows:

$$\beta_Y = (1-\alpha)(1+\chi^L)\frac{Y_t}{L_t}\beta_E^L \tag{1.17}$$

Demonstração. The proof of Lemma 1 comes directly from the definition of the elasticity of hours per worker to the total stock of effective units of labor, χ^L, and the transformation of private formal employment multiplier to a total employment multiplier adjusted for the stock of effective units of labor.

χ^L is formally given by

$$\chi^L = \frac{d_N}{N_t}\frac{L_t}{d_L} = \frac{d_N}{N_t}\frac{FE_t}{d_{FE}}\frac{d_{FE}}{FE_t}\frac{L_t}{d_L} = \frac{\chi}{\Theta^L} \tag{1.18}$$

Using the same steps that were implemented to estimate equation (1.4), we can translate the private formal employment multiplier into a total employment multiplier adjusted for effective units of labor, as follows:

$$\beta_E^L = \Theta^L \beta_{FE}(1/\omega_{fe}^L) \tag{1.19}$$

After substituting equation (1.18) and (1.19) in equation (1.17), we find that

$$\beta_Y = (1-\alpha)(1+\frac{\chi}{\Theta^L})\frac{Y_t}{L_t}\Theta^L\beta_{FE}\frac{1}{\omega_{fe}^L} = (1-\alpha)(\chi+\omega_{fe}^L+\rho\Psi\omega_{if}^L+\eta\omega_{pe}^L)\frac{Y_t}{FE_t}\beta_{FE} \tag{1.20}$$

□

1.8 APPENDIX

Table 1.9: Baseline Regression Results (OLS)

	Cumulative Change in Formal Employment per capita			
	OLS			
	Baseline	Adding Services Empl.	Adding Urbanization Rate	Adding Mobility
	(1)	(2)	(3)	(4)
EA per capita (BRL 100K)	0.130	0.132	0.161	0.312^{***}
	(0.112)	(0.112)	(0.112)	(0.096)
Covid-19 Deaths	0.036	0.086	0.136	0.331^{*}
	(0.122)	(0.123)	(0.123)	(0.178)
Informality Rate	0.001	0.001	-0.005^{**}	-0.008^{**}
	(0.002)	(0.002)	(0.002)	(0.003)
Δ Formal Employment Q12020	-0.430^{*}	-0.432^{**}	-0.430^{*}	-0.279
	(0.222)	(0.220)	(0.220)	(0.233)
Δ Formal Employment 2019	0.008^{***}	0.004^{*}	0.004	0.002
	(0.002)	(0.002)	(0.002)	(0.006)
Share of Services in Employment		-0.011^{***}	-0.008^{***}	0.002
		(0.002)	(0.002)	(0.003)
Urbanization Rate			-0.006	-0.009^{***}
			(0.008)	(0.002)
Overall Mobility				0.024^{***}
				(0.003)
Constant	-0.003^{*}	0.003^{*}	0.009^{***}	0.008^{**}
	(0.002)	(0.002)	(0.002)	(0.003)
States Fixed effects	*Yes*	*Yes*	*Yes*	*Yes*
Observations	5,478	5,478	5,478	2,210
R^2	0.091	0.098	0.102	0.160
Adjusted R^2	0.086	0.093	0.096	0.147
Residual Std. Error	0.013 (df = 5446)	0.013 (df = 5445)	0.013 (df = 5444)	0.011 (df = 2175)
F Statistic	17.626^{***} (df = 31; 5446)	18.508^{***} (df = 32; 5445)	18.660^{***} (df = 33; 5444)	12.164^{***} (df = 34; 2175)

*p<0.1; **p<0.05; ***p<0.01

Note: The unit of analysis is municipalities, the lowest administrative level in Brazil. Data on formal employment creation, and EA transfers refer to the sum over April-September 2020. Google mobility and Covid-19 deaths are measured as the average over April-September 2020. The share of population receiving *Bolsa Familia* benefits is as of December/2019. Services as a share of total employment, urbanization rate, and the informality rate come from the 2010 census.

Table 1.10: Baseline Regressions First Stage Results

	Cumulative Change in EA Transfers per capita (BRL 100K)			
	Baseline	Adding Services Empl.	Adding Urbanization Rate	Adding Mobility
	(1)	(2)	(3)	(4)
Share of BF Receivers in Population	0.039^{***}	0.039^{***}	0.039^{***}	0.032^{***}
	(0.001)	(0.001)	(0.001)	(0.002)
Covid-19 Deaths	0.123^{***}	0.119^{***}	0.106^{***}	0.129^{***}
	(0.016)	(0.016)	(0.016)	(0.028)
Informality Rate	0.003^{***}	0.003^{***}	0.004^{***}	0.005^{***}
	(0.0002)	(0.0002)	(0.0002)	(0.001)
Δ Formal Employment Q12020	-0.003	-0.003	-0.004	-0.005
	(0.004)	(0.004)	(0.004)	(0.008)
Δ Formal Employment 2019	0.001^{***}	0.001^{***}	0.001^{***}	0.001
	(0.0002)	(0.0002)	(0.0002)	(0.001)
Share of Services in Employment		0.001^{***}	0.0003	0.001^{**}
		(0.0002)	(0.0002)	(0.0005)
Urbanization Rate			0.001	0.001^{***}
			(0.001)	(0.0004)
Overall Mobility				-0.004^{***}
				(0.0004)
Constant	0.007^{***}	0.006^{***}	0.005^{***}	0.003^{***}
States Fixed effects	*Yes*	*Yes*	*Yes*	*Yes*
Observations	5,478	5,478	5,478	2,210
R^2	0.704	0.705	0.709	0.587
Adjusted R^2	0.702	0.703	0.707	0.581
Residual Std. Error	0.002 (df = 5446)	0.002 (df = 5445)	0.002 (df = 5444)	0.002 (df = 2175)
F Statistic	417.249^{***} (df = 31; 5446)	406.133^{***} (df = 32; 5445)	400.971^{***} (df = 33; 5444)	91.105^{***} (df = 34; 2175)

*p<0.1; **p<0.05; ***p<0.01

Note: The unit of analysis is municipalities, the lowest administrative level in Brazil. Data on formal employment creation, and EA transfers refer to the sum over April-September 2020. Google mobility and Covid-19 deaths are measured as the average over April-September 2020. The share of Population Receiving the Bolsa Familia as December/2019. Services as a share of total employment, urbanization rate, and the informality rate come from the 2010 census.

CHAPTER 2

DOES THE CREATION OF A SOVEREIGN INFLATION-LINKED BOND MARKET (INFORMATION) MATTER? THE INTERPLAY BETWEEN TERM PREMIA AND THE DEMAND FOR GOVERNMENT SECURITIES

2.1 INTRODUCTION

Public debt bonds carry meaningful information on expected future inflation and real interest rates, associated to economic fundamentals. They are often distorted by various factors associated to current risk perception, market appetite and trends in the external and domestic environment. The main contributions of this paper is the provision of a novel theoretical and empirical analysis on the complementary interplay between a sovereign Inflation Linked Bond (ILB) and a regular fixed rate market. We develop an empirical micro-funded method to extract the term premia and the underling demand for government securities not having directly to rely on level government bonds fundamentals, which are, ex-ante, non-observable variables. To the best of our knowledge, this is the first study that estimates the impact of the opening of an ILB sovereign market on term premia.

We address key questions for sovereign debt market players: i) how to infer more accurately whether domestic public security's prices are effectively reflecting expectations on economic fundamentals?, ii) how the demand for sovereign securities is impacted by the term premia?, iii) how the demand for sovereign fixed rate bond is analytically impacted by the ILB market, and iv) what is the impact of the opening of ILB market on nominal term premia?

To answer the first question we modify the price informativeness model proposed by Dávila and Parlatore (2018) to accommodate government bonds idiosyncrasies showing that term premia metric can be estimated by relative price informativeness taking into account the noise component priced in government yields. To address the second question, we build a simple model to map how the share allocated in risky government bonds of an investor is impacted by the term premia. To response the third question, we applied the fisher decomposition on the optimal allocation derived from our structural model. On the empirical section, the fourth question is depicted using a difference-in-differences approach.

We formally show that government bonds' term premia measures how much an external observer can not learn from prices and, thus, macroeconomic fundamentals. Additionally, we demonstrate that the demand for public debt securities depends on $\frac{1}{\gamma^i}\frac{E[-\Delta\tau^E_{\nu t+1}]}{VAR[-\Delta\tau^E_{\nu t+1}]}$, where γ is the coefficient of relative risk aversion and τ^E_ν is the term premia. Intuitively, the demand for public bonds will be higher if agents expect that the term premia will contract over timer and/or the variation of the term premia is low.

Differently from equities or cooperate bonds, the fundamental of government yields bonds are purely derived from the macroeconomics fundamentals of a given economy. They are expressed by the expectation on the future short term policy rate, inflation, and real interest rates. The fundamental of a government yield can be estimated by a macro-model, a survey of professional forecasters, or models designed to extract fundamentals from yield curves. Cohen et al (2018) found out that distinct models produce different estimates for the levels of fundamentals, but broadly agree on the trends and dynamics. Therefore, the logdifference-stationary environment from Dávila and Parlatore (2018) suits well the purpose of this paper given that the choice of the method of estimation of fundamentals does not become a burden in our exercise.

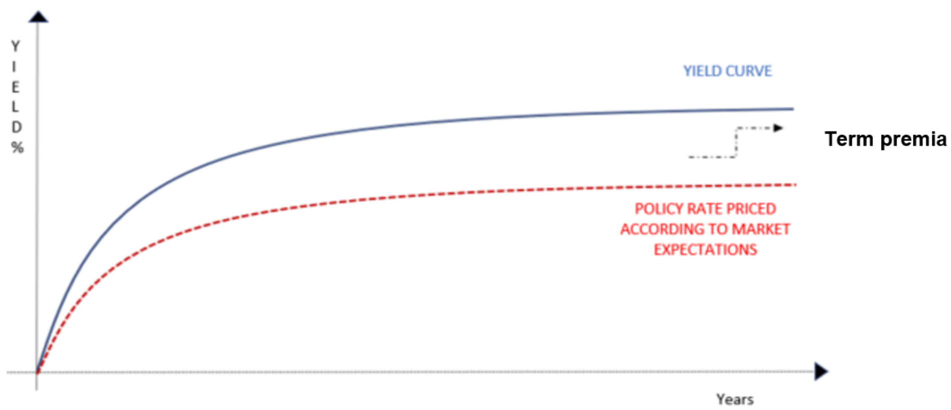


Figure 2.1: Computing the yield curve term premia using professionals' market forecasts to proxy fundamentals

Agents' expectations are a key variable to infer the cost of public debt for a given country. Diebold, Rudebusch and Aruoba (2006) found strong evidence of the effects of macroeconomic variables on movements in the interest curve. In different ways, the literature corroborates this perception through other possibilities: i) hypothesis of expectations of the term structure of interest rates, which establishes that long-term rates reflect the expected path for future short-term rates; ii) the disparity of the effects of monetary policy on the interest rate curve on short and long term rates; and iii) the intertemporal maximization of consumer choices, based on the hypothesis that consumers prefer a stable level to the detriment of fluctuations in income.

As detailed in Section 2.3, our analysis focus in spirit on work of Adriel et al

(2013) as one can read the term premia as a compensation that investors demand for holding the risk that interest rates (fundamentals) may change over the life of the bond. Furthermore, we demonstrate that we could interpret the term premia as a metric of how much an external observer cannot learn about fundamentals from prices.

On the role of ILB for sovereigns, Velandia-Rubiano et al. (2022) find mixed evidence on the cost-effectiveness of ILBs. On one hand, it can be used to lengthen debt maturity and replace FX-linked and FX-denominated securities. On the other hand, introducing ILBs may reduce government securities' trading and liquidity by fragmenting an already small market.

In light of the theoretical insights stemming from price informativeness framework, we conciliate this puzzle by empirically segregating the outcomes of the creation of ILB per se and the dynamics of conventional bonds.By focusing on the opening of ILB market, we discovery that the opening leads to a significant improvement across different term premia metrics for EMs, but it is not significant for AEs.

The paper is structures as follow. Section 2 details our identification strategy to bridge price informativeness from linear regressions to our extension of Dávila and Parlatore (2018) model. Section 3 introduces a fully microfounded model, allowing us to recover model primitives. Section 4 depicts our empirical exercises on the estimation of the impact of the creation of ILB market on nominal term premia and the transmission of inertial nominal term premia shock contingent on existence of an ILB market, while section 5 concludes.

2.2 IDENTIFYING GOVERNMENT BONDS PRICE INFORMATIVENESS

In this section, we introduce the main identification results in the context of a dynamic model with a single government bond yield of a given maturity whose payoff process is stationary.

In this section, we show how to formally identify and estimate price informativeness from an asset pricing equation and a stochastic process for government yield payoffs. In line with Dávila and Parlatore (2018), we derive the main results in the body of the paper in a logdifference-stationary environment.

Model

Time is discrete, with periods denoted by $t = 0{,}1, 2{,}...{,}\infty$. There is a continuum of investors, indexed by $i \in I$, who trade a risky nominal government bond in fixed supply each period at a (log) price p_t. The payoff of the nominal yield in period t+1, x_{t+1}, - as well as the payoff of the real interest rate component, x^r_{t+1}, and the inflation priced in government bonds, x^π_{t+1}- are given by the following stationary AR(1) process:

$$\Delta x^r_{t+1} = \mu_{\Delta x^r} + \rho^r \Delta x^r_t + u^r_t \tag{2.1}$$

$$\Delta x^\pi_{t+1} = \mu_{\Delta x^\pi} + \rho^\pi \Delta x^\pi_t + u^\pi_t \tag{2.2}$$

$$\Delta x_{t+1} = \mu_{\Delta x^r} + \mu_{\Delta x^\pi} + \rho^r \Delta x^r_t + \rho^\pi \Delta x^\pi_t + u^r_t + u^\pi_t \tag{2.3}$$

$$u_t = u^r_t + u^\pi_t \tag{2.4}$$

Where $\mu_{\Delta x^r}$ and $\mu_{\Delta x^\pi}$ are a scalars, $|\rho^r| and |\rho^\pi| < 1$, and where the innovations to the payoff,u_t, have mean zero, finite variance denoted by $Var(u_t) = Var(u^r_t) +$

$2Cov(u_t^r, u_t^\pi) + Var(u_t^\pi) = \sigma_u^2 = \tau_u^{-1}$. It is important to note that u_t re identically and independently distributed over time. The payoff setup is built in such way that the innovation to the t + 1 payoff difference, u_t, is indexed by t — instead of t + 1 — to indicate that investors can potentially learn about the realization of u_t at date t.

We assume that the equilibrium nominal (log) yield difference,Δp_t, the real(log) yield difference,Δp_t^r,and the inflation log difference,Δp_t^π, are given by

$$\Delta p_t = \phi + \phi_0 \Delta x_t^r + \phi_1 \Delta x_t^\pi + \phi_2 \Delta x_{t+1} + \phi_n \Delta n_t \tag{2.5}$$

$$\Delta p_t^r = \xi + \xi_0 \Delta x_t^r + \xi_2 \Delta x_{t+1}^r + \xi_n \Delta n_t^r \tag{2.6}$$

$$\Delta p_t^\pi = \iota + \iota_0 \Delta x_t^\pi + \iota_1 \Delta x_{t+1}^\pi + \iota_n \Delta n_t^\pi \tag{2.7}$$

$$\Delta n_t^r = \mu_{\Delta n^r} + \epsilon_t^{\Delta n^r} \tag{2.8}$$

$$\Delta n_t^\pi = \mu_{\Delta n^\pi} + \epsilon_t^{\Delta n^\pi} \tag{2.9}$$

$$\Delta n_t = n_t^r + n_t^\pi = \mu_{\Delta n} + \epsilon_t^{\Delta n} \tag{2.10}$$

where ι, ι_0, ι_1,$\iota_n, \xi, \xi_0, \xi_2$,$\xi_n, \phi, \phi_0, \phi_1$, and ϕ_n are parameters and where $\Delta n_t, \Delta n_t^r$ and Δn_t^π represents the change in the aggregate component of investors' trading motives on nominal and real yields and inflation that are orthogonal to the yield payoff. Our timing assumes that date t variables, in particular Δx_t and u_t, are realized before the price p_t is determined. Moreover, we assume that u_t and Δn_t are independent.

2.2.1 Government bonds Price Informativeness: Definition

From the perspective of understanding the informativeness of government yields about future payoffs, the key variable of interest is the unbiased signal of the innovation to future payoffs,u_t, incorporated in the price. This endogenous signal, which we denote by ν_t , is given by

$$\nu_t \equiv \frac{\Delta p_t - (\phi + \phi_2\mu_{\Delta x^r} + \phi_2\mu_{\Delta x^\pi} + (\phi_0 + \phi_2\rho^r)x_t^r + (\phi_1 + \phi_2\rho^\pi)x_t^\pi + \phi_n Z_n}{\phi_2} \tag{2.11}$$

where $Z_n = \mu_{\Delta n^r} + \mu_{\Delta n^\pi} + \epsilon_t^{\Delta n^r} + \epsilon_t^{\Delta n^\pi}$

Given equation (2.11), we can write the endogenous unbiased signal about u_t as

$$\nu_t = u_t + \frac{\phi_n}{\phi_2}(\Delta n_t^r - \mu_{\Delta n^r}) + \frac{\phi_n}{\phi_2}(\Delta n_t^r - \mu_{\Delta n^r}) \tag{2.12}$$

One can easily see that the signal, ν_t is unbiased because $E[\nu_t \mid u_t,\Delta x^r, \Delta x^\pi] = u_t$.

Following Dávila and Parlatore (2018) steps', we define two key measures of government yields informativeness: absolute and relative price informativeness. These are the crucial metrics for an external observer who learns about the future payoff from the government yields. Additionally, we define a key metric for government bonds: the relative term premia.

Definition. (Government bonds yield informativeness)

a) Absolute price informativeness, denoted by $\tau_\nu \in [0,\infty)$ is the precision of the unbiased signal about the innovation to the asset payoff contained in the asset price. Given Equation (2.5), it is formally given by

$$\tau_{\nu} \equiv (Var[\nu_t \mid x_{t+1}, \Delta x^r_t, \Delta x^{\pi}_t])^{-1} = (\frac{\phi_2}{\phi_n})^2 \tau_{\Delta n} \tag{2.13}$$

where $\tau_{\Delta n} = (Var[\Delta n_t])^{-1} = (Var[\Delta n^r_t] + 2Cov[\Delta n^r_t, \Delta n^{\pi}_t] + Var[\Delta n^{\pi}_t])^{-1}$

b) Relative price informativeness, denoted by $\nu^R_t \in [0,1]$, is the ratio between absolute price informativeness and the sum of absolute price informativeness and the precision of the innovation to the yield payoff. Given Equation (2.5), it is formally given by

$$\tau^R_{\nu} \equiv \frac{\tau_{\nu}}{\tau_{\nu} + \tau_u} \tag{2.14}$$

where $\tau_u = (Var[u_t])^{-1} = (Var(u^r_t) + 2Cov(u^r_t, u^{\pi}_t) + Var(u^{\pi}_t))^{-1}$

c) Relative term premia, denoted by $\nu^E_t \in [0,1]$, is the ratio between the precision of the innovation to the yield payoff and the sum of absolute price informativeness and the precision of the innovation to the yield payoff. Given Equation (2.5), it is formally given by

$$\tau^E_{\nu} \equiv \frac{\tau_u}{\tau_{\nu} + \tau_u} \tag{2.15}$$

The definition of absolute price informativeness is derived from Blackwell(1953) criterion to rank experiments/signals, which is deeply used in literature on information and learning in financial markets, see, e.g., Vives (2008) and Veldkamp (2011).

Absolute price informativeness reveals, for given realizations of the future and current fundamentals, x_{t+1}, Δx^r_t, and Δx^{π}_t, indicates whether the signal incorporated in the yield is close to the fundamental taking into account the volatility expressed by $\tau_{\Delta n}$. On one hand, absolute price informativeness has a negative relation with the dispersion of the noise,τ_n, and the trading needs given by ϕ_n. On the other hand, τ_ν is positively related with private signals given by ϕ_2. As a result, absolute price informativeness captures how much information about the fundamental can be gained by an uninformed external observer by exclusively observing the yield. When absolute price informativeness is high, an external observer receives a very precise signal about the fundamental by observing the nominal yield Δp_t. On the contrary, when price informativeness is low, Δp_t chiefly represents noise rather than fundamentals.

Relative price informativeness adjusts absolute price informativeness for the precision of the innovation to the fundamental through τ_u. As Dávila and Parlatore (2018) pointed out, this measure expresses how much can be learned by observing the price relative to the volatility of the fundamental. Relative price informativeness measures, for an external observer, how much can be learned from the price relative to the total amount that can be learned. In the body of the paper, we focus on the identification of relative metrics of price informativeness rather than absolute ones because it is easily interpreted and comparable across bonds of different countries and maturities.

Relative term premium is, by construction, the complement of relative price informativeness, providing, thus, a metric about how much an external observer can not learn from the price relative to the total amount that can be learned. Thus, relative term premium can be interpreted as the residual uncertainty about payoff innovations after observing prices. On the one hand, relative term premium definition is similar to the term premia definition highlighted in the literature -such as in Kim and Orphanides (2007) and Cohen et al (2018)- because it gauges a risk compensation for the variance of the expected future short-term policy rate. On the other hand, relative term premia is different from the standard term premia metrics as $\nu_t^E \in [0{,}1]$, and the regular term premia returns a fraction of gross yields.

Price Informativeness: Identification

Proposition 1. (Identifying price informativeness) Let β, β_0, β_1, β_2 denote the coefficients of the following regression of log-price differences on realized and future log-payoff differences:

$$\Delta p_t = \beta + \beta_0 \Delta x_t^r + \beta_1 \Delta x_t^\pi + \beta_2 \Delta x_{t+1} + e_t \tag{R1}$$

where$\Delta p_t = p_t - p_{t-1}$ denotes the date t change in log-yields, $\Delta x_t^r = x_t^r - x_{t-1}^r$, $\Delta x_t^\pi = x_t^\pi - x_{t-1}^\pi$, and $\Delta x_{t+1} = x_{t+1} - x_t$ respectively represent the date t and $t+1$ log-payoff differences of real yields, inflation, and nominal yields, and where $R^2_{\Delta x \Delta x'}$ denotes the Rsquared of Regression R1. Let ζ, ζ_0, ζ_1 indicate the coefficients of the following regression of log-yield differences on realized log-payoff differences:

$$\Delta p_t = \zeta + \zeta_0 \Delta x_t^r + \zeta_1 \Delta x_t^\pi + e_t^\zeta \tag{R2}$$

where $R^2_{\Delta x}$ is the R-squared of Regression R2. Then, relative price informativeness,τ_ν^R, defined in Equation (2.15), can be recovered as

$$\tau_\nu^R = \frac{R^2_{\Delta x \Delta x'} - R^2_{\Delta x'}}{1 - R^2_{\Delta x'}} \tag{2.16}$$

The proof of Proposition 1 relies on identifying the right combination of parameters in the econometric specification defined by Regressions R1 and R2 that maps into the definition of relative price informativeness,τ_ν^R.

PROOF 1. ***(Identifying price informativeness)***. *The R-squareds of both regressions can be expressed as follows*

$$R^2_{\Delta x \Delta x'} = 1 - \frac{Var(e_t)}{Var(\Delta p_t)}$$

$$R^2_{\Delta x'} = \frac{Var(\zeta_0 \Delta x^r_t + \zeta_1 \Delta x^\pi_t)}{Var(\Delta p_t)}$$

After substituting Equation (2.5) in Equation (2.11), the find that

$$\Delta p_t = \phi + \phi_2 \mu_{\Delta x^r} + \phi_2 \mu_{\Delta x^\pi} + \phi_n \mu_{\Delta n^r} + \phi_n \mu_{\Delta n^\pi} + (\phi_0 + \rho^r \phi_2)\Delta x^r_t + (\phi_1 + \rho^\pi \phi_2)\Delta x^\pi_t \\ + \phi_2 u_t + \phi_n(\epsilon^{\Delta n^r}_t + \epsilon^{\Delta n^\pi}_t) \quad (2.17)$$

From the comparison of regression R2 with the structural Equation (2.16), one can see that $\psi = \phi + \phi_2 \mu_{\Delta x^r} + \phi_2 \mu_{\Delta x^\pi} + \phi_n \mu_{\Delta n^r} + \phi_n \mu_{\Delta n^\pi}$, $\psi_0 = \phi_0 + \rho^r \phi_2$, $\psi_1 = \phi_1 + \rho^\pi \phi_2$ *and* $\epsilon^\psi_t = \phi_2 u_t + \phi_n(\epsilon^{\Delta n^r}_t + \epsilon^{\Delta n^\pi}_t)$*. By comparing Regression R1 with the structural Equation (2.5), it follows that* $\beta = \phi$, $\beta_0 = \phi_0$, $\beta_1 = \phi_1$, $\beta_2 = \phi_2$, *and* $e_t = \phi_n \Delta_{n_t}$.

From Equation (2.16), the following variance decomposition must hold

$$\begin{aligned} Var[\Delta p] &= Var[\zeta_0 \Delta x^r_t + \zeta_1 \Delta x^\pi_t] + Var[\phi_2 u_t + \phi_n(\epsilon^{\Delta n^r}_t + \epsilon^{\Delta n^\pi}_t)] \\ &= Var[\zeta_0 \Delta x^r_t + \zeta_1 \Delta x^\pi_t] + \phi_2^2 Var[u_t] + Var[e_t] \end{aligned}$$

which can be rearranged to express $\frac{\tau_\nu}{\tau_u}$ *as follows*

$$1 = \frac{Var[\zeta_0 \Delta x^r_t + \zeta_1 \Delta x^\pi_t]}{Var[\Delta p]} + \frac{Var[e_t]}{Var[\Delta p]}\left(\frac{\phi_2^2}{Var[e_t]} Var[u_t] + 1\right)$$

$$1 = R^2_{\Delta x'} + (1 - R^2_{\Delta x \Delta x'})\left(\frac{\tau_\nu}{\tau_u} + 1\right) \implies \frac{\tau_\nu}{\tau_u} = \frac{R^2_{\Delta x \Delta x'} - R^2_{\Delta x'}}{1 - R^2_{\Delta x'}}$$

As a result, relative price informativeness can be written as

$$\tau_\nu^R = \frac{\tau_\nu}{\tau_\nu + \tau_u} = \frac{1}{1 + \frac{1}{\frac{\tau_\nu}{\tau_u}}} = \frac{R^2_{\Delta x \Delta x'} - R^2_{\Delta x'}}{1 - R^2_{\Delta x'}}$$

$\square$

Numerically, a relative price informativeness of, for instance, 0.6, means that the initial uncertainty of an external observer about the innovation to the future payoff is reduced by 60 percent after learning from the price — this reading follows from the fact that $(\tau_\nu + \tau_u)^{-1} = \frac{1-\tau_\nu^R}{\tau_u}$.

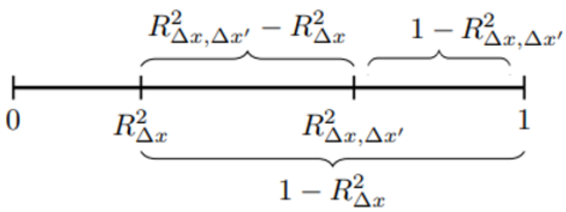


Figure 2.2: Explaining relative price informativeness

Geometrically, Figure 1 provides another angle to interpret Equation (2.14). The denominator $1 - R^2_{\Delta x'}$ can be interpreted as the residual uncertainty about future payoffs after conditioning on the realized date t yield payoff. In turn, the numerator, $R^2_{\Delta x \Delta x'}$, can be understood as the percentage reduction in uncertainty about future payoffs after observing the asset price at date t in addition to the realized payoff.

Proposition 2. (Relative term premium). Relative term premium ,τ_ν^E, defined in Equation (2.15), can be recovered as

$$\tau_\nu^E = \frac{1 - R^2_{\Delta x \Delta x'}}{1 - R^2_{\Delta x'}} \tag{2.18}$$

PROOF 2. ***(Relative term premium)***. *By definition the relative term premium is the complement of relative price informativeness. Therefore, the proof of proposition 2 is straightforward as* $R^2_{\Delta x \Delta x'} - R^2_{\Delta x'} + 1 - R^2_{\Delta x \Delta x'} = 1 - R^2_{\Delta x'}$.

□

Proposition 3. (Identifying absolute price informativeness) Let β, β_0, β_1, β_2 denote the coefficients of the following regression of nominal log-price differences on realized and future log-payoff differences described by (R1). Let β', β'_0, β'_2 represent the coefficients of the following regression of real log-price differences on realized and future log-payoff differences described by (R1'). Finally, let β'', β''_1, β''_2 denote the coefficients of the following regression of inflation log-price differences on realized and future log-payoff differences described by (R1").Then, we can recover absolute price informativeness of a nominal government yield,τ_ν, real interest rates,τ^r_ν,and inflation,τ^π_ν, as follows

$$\Delta p^r_t = \beta' + \beta_{0'} \Delta x^r_t + \beta_{2'} \Delta x^r_{t+1} + e^r_t \tag{R1'}$$

$$\Delta p^\pi_t = \beta'' + \beta_{1''} \Delta x^\pi_t + \beta_{2''} \Delta x^\pi_{t+1} + e^\pi_t \tag{R1''}$$

$$\tau_\nu = \frac{\beta_2^2}{\sigma_e^2} = \frac{\phi_2^2}{\phi_n^2 \tau_{\Delta_n}^{-1}} = (\frac{\phi_2}{\phi_n})^2 \tau_{\Delta_n} \tag{2.19}$$

$$\tau_{\nu^r} = \frac{\beta_{2'}^2}{\sigma_{e^r}^2} = \frac{\xi_2^2}{\xi_{n^r}^2 \tau_{\Delta_n^r}^{-1}} = (\frac{\xi_2}{\xi_n^r})^2 \tau_{\Delta_n^r} \tag{2.20}$$

$$\tau_{\nu^\pi} = \frac{\beta_{2''}^2}{\sigma_{e^\pi}^2} = \frac{\iota_2^2}{\iota_{n^\pi}^2 \tau_{\Delta_n^\pi}^{-1}} = (\frac{\iota_2}{\iota_n^r})^2 \tau_{\Delta_n^\pi} \tag{2.21}$$

PROOF 3. ***(Identifying absolute price informativeness)***.*The proof of proposition 3 comes directly from the comparison of Regression R1 with the structural Equation (2.5), which implies that* $\beta = \phi$, $\beta_0 = \phi_0$, $\beta_1 = \phi_1$, $\beta_2 = \phi_2$, *and* $e_t = \phi_n \epsilon_t^{\Delta_{n_t}}$ *and, thus,* $\sigma_e^2 = Var[e_t] = \phi_n^2 Var[\epsilon_t^{\Delta_{n_t}}] = \phi_n^2 {\tau_{\Delta n}}^{-1}$. *The comparison* *of Regression R1' with the Equation (2.6) yields that* $\beta' = \xi$, $\beta_0' = \xi_0$, $\beta_2' = \xi_2$, *and* $e_t^r = \xi_{n^r}\epsilon_t^{\Delta_{n_t}r}$ *and, thus,* $\sigma_{e^r}^2 = Var[e_t^r] = \xi_{n^r}^2 Var[\epsilon_t^{\Delta_{n_t^r}}] = \xi_{n^r}^2\tau_{\Delta n^r}^{-1}$ *. Lastly, the matching of Regression R1" with the Equation (2.7) implies that* $\beta'' = \iota$, $\beta_0' = \iota_0$, $\beta_2' = \iota_1$, *and* $e_t^r = \iota_{n^\pi}\epsilon_t^{\Delta_{n_t}\pi}$ *and, as a result,* $\sigma_{e^\pi}^2 = Var[e_t^\pi] = \iota_{n^\pi}^2 Var[\epsilon_t^{\Delta_{n_t^\pi}}] = \xi_{n^\pi}^2\tau_{\Delta n^\pi}^{-1}$.

□

2.3 STRUCTURAL MODEL

Environment We consider a tractable overlapping generations model. Time is discrete, with dates denoted by $t = 0, 1, 2,...,\infty$. The economy is populated by a continuum of investors, indexed by $i \in I$, who live for two dates. Each investor I is born with wealth w_0^i and has well-behaved expected utility preferences over his terminal wealth w_1^i, with flow utility given by $U_i(w_1^i)$, where $U_i'(.) > 0$ and $U_i''(.) < 0$. We assume that the distribution of initial wealth is bounded and i.i.d. across time and investor types.

There are two types of assets in the economy: a short-term risk-free asset in perfectly elastic supply, with gross return $R^f > 1$, and a risky government bond asset with a given maturity m in fixed supply Q. The government bond price valuated by fundamentals P^X depend on the fundamental yield Y^X, which denotes the expected average short-term policy rate from date $t + m$ according to fundamentals with the (log) payoff is $x_t = ln(Y_t^X)$. The government bond price traded is denoted by P with an underlying yield Y and (log) yield $p_t = ln(Y_t)$. The law of motion of the fundamentals x_t, x_t^r, and x_t^π are given by

$$\Delta x_{t+1}^r = \mu_{\Delta x^r} + u_t^r \tag{2.22}$$

$$\Delta x_{t+1}^{\pi} = \mu_{\Delta x^{\pi}} + u_t^{\pi} \tag{2.23}$$

$$\Delta x_{t+1} = \mu_{\Delta x^{r}} + \mu_{\Delta x^{\pi}} + u_t \tag{2.24}$$

where $u_t = u_t^r + u_t^{\pi}$,$\Delta x_{t+1}^r = x_{t+1}^r - x_t^r$, $\Delta x_{t+1}^{\pi} = x_{t+1}^{\pi} - x_t^{\pi}$,$\mu_{\Delta x^r}$ and $\mu_{\Delta x^{\pi}}$ are a scalars,$x_0^r = 0$, and $x_0^{\pi} = 0$. The realized payoff x_t^r and x_t^{π} are common knowledge to all investors before the yield p_t is determined. The realized payoff at date $t+1$, x_{t+1}, is only revealed to investors at date $t+1$. It is important to highlight that equation (2.23) is a special case of equation (2.5) when $\rho^r = 0$ and $\rho^{\pi} = 0$.

We assume that investors receive private signals about the innovation to the risky asset payoff. Formally, each investor receives a signal about the payoff innovation u_t given by

$$s_t^i = u_t + \epsilon_{st}^i$$

with

$$\epsilon_{st}^i \sim N(0,\tau_s^{-1})$$

where $\epsilon_{st}^i \perp \epsilon_{st}^j$ for all $i \neq j$ and $u_t \perp \epsilon_{st}^i$ for all t and for all i.

We also assume that investors have additional private trading motives coming from random heterogeneous priors that are random in the aggregate. Formally, each investor i born at date t has a prior over the innovations to the payoff difference u_t given by

$$u_t \sim i,tN(\overline{n}_t^i,\tau_u^{-1})$$

where

$$\overline{n}_t^i = n_t + \epsilon_{\overline{n}_t}^i$$

with

$$\epsilon^i_{\overline{n}_t} N(0,\tau^{-1}_{\overline{n}})$$

and

$$\Delta n_t = \mu^r_{\Delta n^r} + \mu^\pi_{\Delta n^\pi} + \epsilon^{\Delta n^r}_t + \epsilon^{\Delta n^\pi}_t = \mu_{\Delta n} + \epsilon^{\Delta n}_t$$

with

$$\epsilon^{\Delta n}_t \sim i,tN(0,\tau^{-1}_n)$$

where $n_0 = 0$, $\mu_{\Delta n}$ is a scalar, and $\epsilon^{\Delta n}_t \perp \epsilon^i_{\overline{n}_t}$ for all i and for all t. One can read he variable n_t as the aggregate sentiment in the economy. n_t is not observed and acts as a source of aggregate noise, preventing the yield from being fully revealing. In addition, we assume that $u_{t+s} \sim ,i,tN(0,\tau^{-1}_u)$ for all $s > 0$.

Each investor i born at date t optimally chooses a portfolio share in the risky asset, denoted by θ^i_t, to solve

$$\max_{\theta^i_t} E^i_t[U_i(w^i_1)] \tag{2.25}$$

subject to a wealth accumulation constrain

$$w^i_1 = (R^f + \theta^i_t(\frac{\frac{P_{t+1}}{P^X_{t+1}}}{\frac{P_t}{P^X_t}}))w^i_0 = (R^f + \theta^i_t(\frac{P_{t+1}}{P_t}[\frac{P^X_{t+1}}{P^X_t}]^{-1}))w^i_0 \tag{2.26}$$

where the information set of an investor i in period t is given by

$$I^i_t = \{s^i_t,\overline{n}^i_t,\{X_s\}_{s\leq t},\{P_s\}_{s\leq t}\}$$

An investor i faces a trade-off between allocating its resources at the short-term policy rate, R^f, or at a government bond with a maturity m, whose return is determined by the term premium over investor i expectation on future short-term policy rate according to fundamentals which denotes an extra return (a risk premium) that investors demand to compensate them for the risk associated with a non-short term bond.

The economic return of a government bond The wealth accumulation constrain given by equation (2.26) entails the profitability of government bond is depicted by $\frac{PV_{t+1}}{PV_t}$ where $PV_{t+1} = \frac{P_{t+1}}{P_{t+1}^X}$ and $PV_t = \frac{P_t}{P_t^X}$ are the price-to-value ratio of periods $t+1$ and t respectively. Intuitively, an overvaluation in period $t+1$ (higher PV_{t+1}) in relation to fundamentals increases the bond return all else equal. Conversely, an undervaluation in period t (lower PV_t) implies a greater profitability. A key implication of this setup is that investors are seeking arbitrages opportunities derived from the spread between market prices and fundamentals of a bond with maturity m taking into account the opportunity cost given by risk-free rate.

Definition. (Equilibrium) A stationary rational expectations equilibrium in linear strategies is a set of portfolio shares θ_t^i for each investor i at date t and a price function P_t such that: i) θ_t^i maximizes the investor i's expected utility given his information set and ii) the price function P_t is such that the market for the risky asset clears at each date t, that is, $\int \theta_t^i w_0^i \, di = Q$

Even though for this type of models it is not possible to characterize in closed-form the portfolio problem solved by investors and the equilibrium price, Dávila and Parlatore (2018) showed that it is possible to find a closed-form solution to the model in approximate form.

The optimally condition of an investor who maximizes Equation (2.25) subject to the wealth accumulation constraint in Equation (2.26) is given by

$$E[U'(w_1^i)(\frac{\frac{P_{t+1}}{P_{t+1}^X}}{\frac{P_t}{P_t^X}}) \mid I_t^i] = 0 \tag{2.27}$$

Similar to Dávila and Parlatore (2018), we approximate the first-order condition expressed in equation (2.27) in five steps. First, we use a first-order Taylor expansion of an investor's future marginal utility $U'(w_1^i)$ around the current date wealth level w_0^i. As a result, we approximate $U'(w_1^i)$ as follow

$$U'(w_1^i) \approx U'(w_0^i) + U''(w_0^i)\Delta w_1^i \tag{2.28}$$

which permit us to write equation (2.27) as

$$U'(w_0^i)E[\frac{\frac{P_{t+1}}{P_{t+1}^X}}{\frac{P_t}{P_t^X}}] + U''(w_0^i)w_0^iE[((R^f - 1) + \theta_t^i(\frac{\frac{P_{t+1}}{P_{t+1}^X}}{\frac{P_t}{P_t^X}}))(\frac{\frac{P_{t+1}}{P_{t+1}^X}}{\frac{P_t}{P_t^X}})] \approx 0$$

Second, we impose that terms that involve the product of two or more net interest rates are negligible. In continuous time, these terms would be of order $(dt)^2$, implying that

$$((R^f - 1)E[\frac{\frac{P_{t+1}}{P_{t+1}^X}}{\frac{P_t}{P_t^X}}] \approx 0$$

and

$$(E[\frac{\frac{P_{t+1}}{P_{t+1}^X}}{\frac{P_t}{P_t^X}}])^2 \approx 0$$

which allow us to recovery equation (2.27) as follow

$$U'(w_0^i)E[\frac{\frac{P_{t+1}}{P_{t+1}^X}}{\frac{P_t}{P_t^X}}] + U''(w_0^i)w_0^i\theta_t^iVAR[\frac{\frac{P_{t+1}}{P_{t+1}^X}}{\frac{P_t}{P_t^X}}] \approx 0$$

Thus, we can write an investor's risky portfolio share θ^i_t as

$$\theta^i_t \approx \frac{1}{\gamma^i} \frac{E[\frac{\frac{P_{t+1}}{P^X_{t+1}}}{\frac{P_t}{P^X_t}}]}{VAR[\frac{\frac{P_{t+1}}{P^X_{t+1}}}{\frac{P_t}{P^X_t}}]} \tag{2.29}$$

where $\gamma^i \equiv -\frac{w^i_0 U''(w^i_0)}{U'(w^i_0)}$ indicates the coefficient of relative risk aversion. These coefficients are time invariant since we have assumed that the distribution of investor types is time invariant and the wealth distribution across time and investor type is i.i.d.

Third, we write the bond prices P_{t+1}, P_t, P^X_{t+1}, and P^X_t as a perpetuity as follow

$$\frac{\frac{P_{t+1}}{P^X_{t+1}}}{\frac{P_t}{P^X_t}} = \frac{Y^X_{t+1}}{Y_{t+1}} \frac{Y_t}{Y^X_t}$$

Fourth, we take a log-linear approximation of the yields around the predetermined term premium variation. Formally, we find that

$$ln(\frac{Y^X_{t+1}}{Y_{t+1}} \frac{Y_t}{Y^X_t}) = \Delta x_{t+1} - \Delta p_{t+1}$$

Fifth, according to the traditional definition of term premia (the change in yields not explained by the underlying changes in fundamentals), we can interpret $\Delta p_{t+1} - \Delta x_{t+1}$ as the change in the term premia. Accordingly, relative term premia definition, $\tau^E_{\nu t+1}$, gauges this differential in narrower sense as the residual uncertainty about fundamentals innovations after observing prices. Nonetheless, we use the stylized fact highlighted by Cohen et al (2018) of the similarity of term premia trends to do the following approximation $\Delta p_{t+1} - \Delta x_{t+1} \approx \Delta \tau^E_{\nu\, t+1}$. Hence, we can express an investor's risky portfolio share, θ^i_t, as follow

$$\theta_t^i \approx \frac{1}{\gamma^i}\frac{E[\Delta x_{t+1} - \Delta p_{t+1}]}{VAR[\Delta x_{t+1} - \Delta p_{t+1}]} \approx \frac{1}{\gamma^i}\frac{E[-\Delta \tau_{\nu\ t+1}^E]}{VAR[-\Delta \tau_{\nu\ t+1}^E]} \tag{2.30}$$

Forming expectations In order to characterize the equilibrium it is necessary to characterize investors' expectations. We conjecture and subsequently verify that $E[\Delta x_{t+1} - \Delta p_{t+1}]$ is linear in s_t^i, $\overline{n}_t^i$,x_t^r , and x_t^π and $VAR[\Delta x_{t+1} - \Delta p_{t+1}]$ is constant. Assuming this conjecture, θ_t^i is linear function of s_t^i, $\overline{n}_t^i$,x_t^r , and x_t^π that can be written as follow

$$\theta_t^i \approx \alpha_s^i s_t^i + \alpha_{x^r}^i x_t^r + \alpha_{x^\pi}^i x_t^\pi + \alpha_n^i \overline{n}_t^i - \alpha_p^i p^t + \psi^i \tag{2.31}$$

These coefficients are time invariant given that we have assumed that the distribution of investor types is time invariant and the wealth distribution across time and investor type is i.i.d,. This expression and the market clearing conditions $\int \theta_t^i w_0^i,di = Q$ yield that

$$p^t = \frac{\overline{\alpha_s}}{\overline{\alpha_p}}u_t + \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}}x_t^\pi + \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}}x_t^r + \frac{\overline{\alpha_n}}{\overline{\alpha_p}}\overline{n}_t + \frac{\overline{\psi}}{\overline{\alpha_p}} \tag{2.32}$$

where $\overline{\alpha}_h \equiv \int \alpha_h^i w_0^i,di$ for $h = \{x^r, x^\pi, n, p\}$ and $\overline{\psi} = \int \psi^i w_0^i,di - Q$. As in Dávila and Parlatore (2018) and Vives (2008), we make use of the Strong Law of Large Numbers, since the sequence of independent random variables $\{\alpha_s^i w_0^i \epsilon_{st}^i,\alpha_n^i w_0^i \epsilon_{\overline{nt}}^i\}$ has uniformly bounded variance and mean zero, allowing us to rewritten equation (2.32) as follow

$$p^t = \frac{\overline{\alpha_s}}{\overline{\alpha_p}}u_t + (\frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} - \frac{\overline{\alpha_s}}{\overline{\alpha_p}})x_t^\pi + (\frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} - \frac{\overline{\alpha_s}}{\overline{\alpha_p}})x_t^r + \frac{\overline{\alpha_s}}{\overline{\alpha_p}}x_{t+1} + \frac{\overline{\alpha_n}}{\overline{\alpha_p}}n_t + \frac{\overline{\psi}}{\overline{\alpha_p}} - \frac{\overline{\alpha_s}}{\overline{\alpha_p}}\mu_{\Delta x} \tag{2.33}$$

In this model, investors extract information from government bonds yields. The information incorporated in the yield for an investor in the model is

$$\hat{\nu} = \frac{\overline{\alpha_p}}{\overline{\alpha_s}}(p_t - (\frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}}x_t^\pi + \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}}x_t^r + \frac{\overline{\alpha_n}}{\overline{\alpha_p}}\mu_{\Delta x} + \frac{\overline{\psi}}{\overline{\alpha_p}})) \tag{2.34}$$

which has a precision $\tau_{\hat{\nu}} \equiv VAR[\hat{\nu} \mid u_t, \{x_s\}_{s\leq t}, p_{t-1}]^{-1} = (\frac{\overline{\alpha_s}}{\overline{\alpha_n}})^2\tau_{\Delta n}$ .It is important to highlight that we express by ν the unbiased signal of u_t incorporated in the change in log yields Δp_t and by $\hat{\nu}$ the unbiased signal about u_t contained in the log yield p_t. Considering the investor's information set, $VAR[n_t \mid u_t, \{x_s\}_{s\leq t}, p_{t-1}] = VAR[\Delta n_t \mid u_t, \{x_s\}_{s\leq t}, p_{t-1}]$. Therefore, we have that

$$E_t^i[u_t] = E[u_t \mid I_t^i] = \frac{\tau_s s_t^i + \tau_u \overline{n}_t^i + \tau_{\hat{\nu}}\hat{\nu}}{\tau_s + \tau_u + \tau_{\hat{\nu}}}$$

$$E[u_t \mid I_t^i] = \frac{\tau_s s_t^i + \tau_u \overline{n}_t^i + \tau_{\hat{\nu}}(\frac{\overline{\alpha_p}}{\overline{\alpha_s}}(p_t - (\frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}}x_t^\pi + \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}}x_t^r + \frac{\overline{\alpha_n}}{\overline{\alpha_p}}\mu_{\Delta x} + \frac{\overline{\psi}}{\overline{\alpha_p}})))}{\tau_s + \tau_u + \tau_{\hat{\nu}}}$$

$$VAR[u_t \mid I_t^i] = (\tau_s + \tau_u + \tau_{\hat{\nu}})^{-1}$$

These expressions means that our conjecture about θ_t^i is satisfied. To see this, it is key to note that

$$E[\Delta x_{t+1} - \Delta p_{t+1}] = \mu_{\Delta x} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}}\mu_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}}\mu_{x^\pi} - \frac{\overline{\alpha_s}}{\overline{\alpha_p}}\mu_{\Delta x} + (1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}}w_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}}w_{x^\pi})E[u_t]$$

where we considered that $w_{x^r} = \frac{E[u_t^r]}{E[u_t]}$, $w_{x^\pi} = \frac{E[u_t^\pi]}{E[u_t]}$ are scalars. We used that $E[u_{t+1}] = 0$, and $E[\epsilon_{t+1}^{\Delta n}] = 0$. In addition, we have that

$$VAR[\Delta x_{t+1} - \Delta p_{t+1}] = (1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} w_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} w_{x^\pi})^2 VAR[u_t]$$

$$= (1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} w_{tx^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} w_{tx^\pi})^2 (\tau_s + \tau_u + \tau_{\hat{\nu}})^{-1}$$

Using these expressions in the first-order condition and matching coefficients gives

$$\alpha^i_{x^r} = \frac{1}{\gamma^i} \frac{-\tau_{\hat{\nu}} \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_s}}}{(1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} w_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} w_{x^\pi})}$$

$$\alpha^i_{x^\pi} = \frac{1}{\gamma^i} \frac{-\tau_{\hat{\nu}} \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_s}}}{(1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} w_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} w_{x^\pi})}$$

$$\alpha^i_s = \frac{1}{\gamma^i} \frac{\tau_s}{(1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} w_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} w_{x^\pi})}$$

$$\alpha^i_n = \frac{1}{\gamma^i} \frac{\tau_u}{(1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} w_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} w_{x^\pi})}$$

$$\alpha^i_p = \frac{1}{\gamma^i} \frac{\tau_{\hat{\nu}} \frac{\overline{\alpha_p}}{\overline{\alpha_s}}}{(1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} w_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} w_{x^\pi})}$$

$$\psi^i = \frac{1}{\gamma^i} \frac{\mu_{\Delta x} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} \mu_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} \mu_{x^\pi} - \frac{\overline{\alpha_s}}{\overline{\alpha_p}} \mu_{\Delta x} + (1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} w_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} w_{x^\pi}) - \tau_{\hat{\nu}} (\frac{\overline{\alpha_n}}{\overline{\alpha_s}} + \overline{\psi})}{(1 + \frac{\overline{\alpha_s}}{\overline{\alpha_p}} - \frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} w_{x^r} - \frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} w_{x^\pi})^2}$$

Lemma 1 The change in yields described in Equation (2.5) in the identification framework in Section 2 can be calculated endogenously as an approximation of the equilibrium price process in the model described in this section, i.e., the equilibrium price process is given by

$$\Delta p_t \approx \phi + \phi_0 \Delta x_t^r + \phi_1 \Delta x_t^\pi + \phi_2 \Delta x_{t+1} + \phi_n \Delta n_t$$

where the coefficients $\phi = 0$, $\phi_0 = (\frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} - \frac{\overline{\alpha_s}}{\overline{\alpha_p}})$, $\phi_1 = (\frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} - \frac{\overline{\alpha_s}}{\overline{\alpha_p}})$, $\phi_2 = \frac{\overline{\alpha_s}}{\overline{\alpha_p}}$, and $\phi_n = \frac{\overline{\alpha_n}}{\overline{\alpha_p}}$.

PROOF 4. ***Lemma 1*** *Taking the differences of equation (2.32) in $t+1$ and t implies the following*

$$\Delta p^t = (\frac{\overline{\alpha_{x^\pi}}}{\overline{\alpha_p}} - \frac{\overline{\alpha_s}}{\overline{\alpha_p}})\Delta x_t^\pi + (\frac{\overline{\alpha_{x^r}}}{\overline{\alpha_p}} - \frac{\overline{\alpha_s}}{\overline{\alpha_p}})\Delta x_t^r + \frac{\overline{\alpha_s}}{\overline{\alpha_p}}\Delta x_{t+1} + \frac{\overline{\alpha_n}}{\overline{\alpha_p}}\Delta n_t$$

Then, the matching process to map the coefficients comes directly from the comparison with equation (2.5).

□

2.3.1 Fisher´s decomposition of investors' optimal allocation in sovereign bonds

In this section, we analytically demonstrate that the demand for fixed rate bonds is positively impacted by the demand/ information of inflation linked bonds. We argue that well-designed and deep ILB market unveils information on both real interest rates and inflation (break- even inflation), enhancing the yield informativeness, reducing both real interest rate and inflation term premia and, thus, boosting the demand for conventional nominal bonds.

Departing from equation (2.30), we can express the share of an investor's,i, portfolio allocated in risky nominal sovereign fixed-rate bond,$\theta_t^{i,nominal}$, as follow

$$\theta_t^{i,nominal} \approx \underbrace{\frac{1}{\gamma^i}\frac{E[-\Delta\tau_{\nu t+1}^{E,nominal}]}{VAR[-\Delta\tau_{\nu t+1}^{E,nominal}]}}_{\text{demand for fixed rate bonds}} \approx \underbrace{\theta_t^{i,real}\omega_{real} + \theta_t^{i,inflation}\omega_{inflation}}_{\text{Fisher decomposition}} \tag{2.35}$$

where we define $\theta_t^{i,real}$, $\theta_t^{i,inflation}$, ω_{real}, $\omega_{inflation}$ as follow

$$\theta_t^{i,real} \approx \frac{1}{\gamma^i}\frac{E[-\Delta\tau_{\nu t+1}^{E,real}]}{VAR[-\Delta\tau_{\nu t+1}^{E,real}]} \tag{2.36}$$

$$\theta_t^{i,inflation} \approx \frac{1}{\gamma^i}\frac{E[-\Delta\tau_{\nu t+1}^{E,inflation}]}{VAR[-\Delta\tau_{\nu t+1}^{E,inflation}]} \tag{2.37}$$

$$\omega_{real} = \frac{VAR[-\Delta\tau_{\nu t+1}^{E,real}]}{VAR[-\Delta\tau_{\nu t+1}^{E,nominal}]} \tag{2.38}$$

$$\omega_{inflation} = \frac{VAR[-\Delta\tau_{\nu t+1}^{E,inflation}]}{VAR[-\Delta\tau_{\nu t+1}^{E,nominal}]} \tag{2.39}$$

From equation (2.35), one can easily see that the demand for fixed-rate bonds, $\theta_t^{i,nominal}$, is positively related with the demand for ILB, $\theta_t^{i,real}$, and break-even inflation, $\theta_t^{i,inflation}$.

We can also rewrite equation (2.35) to express the demand for ILBs, $\theta_t^{i,real}$, as follow

$$\theta_t^{i,real} \approx \theta_t^{i,nominal}\omega_{real}^{-1} - \theta_t^{i,inflation}\frac{\omega_{inflation}}{\omega_{real}} \tag{2.40}$$

Intuitively, equation (2.40) shows that the demand for ILBs is negatively related to the demand for break-even inflation. As such, an environment of deteriorating inflation expectations (an increase of the inflation term premia) has different effects on ILBs and conventional bonds. The demand for ILBs would be positively impacted while the allocation on conventional bonds would be reduced other things equal.

2.4 EMPIRICAL ANALYSES

In this section, we make use of the insights of the structural model detailed in Section 2.3 to construct and analyze a shock triggered by the opening of ILB sovereign market.

2.4.1 The opening of an ILB market

Data To track the opening of a sovereign ILB market we relied on the dataset of General Government Debt composition by instrument (fixed-rate, floaters, ILBs, and FX) provided by Baking of International Settings (BIS), covering 30 countries on yearly basis from 1995 to 2021. We set a cut-off for ILB of 5 percent[1] of total stock of the General Government debt to define the opening of a sovereign ILB market. As the bulk of the opening happened in late nineties and early 2000, we do not have data on local bond secondary market for several countries. As a result, we used alternatives term premia measures such as the sovereign credit rating (average of Fitch, Moodys, and SP rating) and the 5Y CDS depicted by the work of Kose et al. (2022) that built a broad cross-country database of 30 fiscal space

[1] We also run robustness tests using differents cut-offs of 3,2, and 1 percent that are broadly in line with the results shown in Figures 3, 4 and 5.

metrics over the period of 1990-2021. Furthermore, we use this database to map the developments of local sovereign market underpinned by the creation of ILBs structure focusing on changes of General Government debt in foreign currency and debt average maturity.

Table 2.1: Summary Statistics of EMs treated units

Statistic	N	Mean	St. Dev.	Min	Median	Max
ILB as a share of GGGD	8	24.86	25.28	1.41	21.76	78.91
Year of ILB market opening	8	2003	7	1995	2003	2015
Average of credit agencies sovereing rating (1-21 scale)	8	11.71	3.06	5.82	11.97	15.77
CDS5Y (bps)	8	594	1,106	80	205	3,314
GGGD in foreign currency (percent of total)	5	25.60	14.02	5.18	29.43	42.47
Sovereign debt average maturity (years)	8	12.47	3.49	6.19	12.71	16.74
GGGD (share of GDP)	8	47.45	18.60	16.32	46.32	71.45
Primary balance (share of GDP)	8	0.14	0.84	−1.32	0.53	1.04

Note: The unit of analysis is EMs sovereigns, which are classified using IMF´s Fiscal Monitor taxonomy, aggregated in a yearly basis. Data on ILB as a share of GGGD (average centered at the country level) and year of ILB market opening come from the BIS spaning from 1995 to 2021. The remainder of the variables comes from averages centered at the country level using the database built by Kose et al. (2022) over the period of 1990-2021.

Table 2.2: Summary Statistics of EMs never-treated units

Statistic	N	Mean	St. Dev.	Min	Median	Max
Average of credit agencies sovereing rating (1-21 scale)	46	9.63	2.81	5.44	8.87	15.84
CDS5Y (bps)	17	306	285	118	223	1,176
GGGD in foreign currency (percent of total)	16	41	25.00	0.00	43	77
Sovereign debt average maturity (years)	33	8.95	4.26	3.21	7.24	19.50
GGGD (share of GDP)	51	48.37	23.88	13.49	41.21	118.34
Primary balance (share of GDP)	50	−0.28	2.33	−6.09	−0.53	7.21

Note: The unit of analysis is EMs sovereigns, which are classified using IMF´s Fiscal Monitor taxonomy, aggregated in a yearly basis. The variables comes from averages centered at the country level using the database built by Kose et al. (2022) over the period of 1990-2021.

Empirical strategy of the event-study We follow the approach of Callaway and Sant'Anna (2021) as traditional Two-Way Fixed Effects (TWFE) event-study may be significantly biased. There is a growing literature on the casual interpretations of TWFE regressions[2] pointing out that TWFE linear regression should not

[2]For a detailed discussion see Athey and Imbens (2018; Borusyak and Jaravel (2017); Callaway and Sant'Anna (2021); de Chaisemartin and D'Haultfoeuille (2020); Goodman-Bacon (2018); Sun and Abraham (2021)

be used to highlight treatment effect dynamics. Callaway and Sant'Anna (2021) does not rely on restricting treatment effect heterogeneity assuming only the no-anticipation condition and the existence of parallel trend based on a never-treated and/or not-yet-treated group.

Results The creation of a sovereign Inflation-Linked Bond (ILB) market leads to a significant improvement across different term premia metrics for EMs, but it is not significant for AEs. We find that EMs use ILB market both to migrate from external to local debt and to expand its debt maturity. On the mechanisms, in an environment of deteriorated expectations on Inflation term premia, the demand for ILB is higher than for conventional nominal bonds. As such, ILBs are well-positioned to unlock the migration both from local FX-linked and short fixed-rate debt. Stylized facts shown by Velandia-Rubiano et al. (2022) point out that ILBs were indeed used as a tool to induce an improvement of debt profile by some EMs, such as as Brazil, Chile, and Israel.

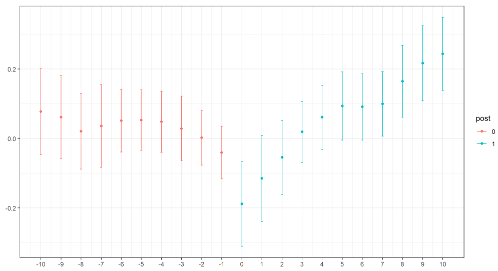
Analytically, recovering our simple theoretical model, equation (2.40) details that the demand for ILBs is negatively related to the demand for break-even inflation. As a result, an environment of deteriorating inflation expectations (an increase of the inflation term premia) has different effects on ILBs and conventional bonds. The demand for ILBs would be positively impacted while the allocation on conventional bonds would be reduced other things equal. Considering that the FX pass-through is effectively operating, ILBs could also be a substitute asset to local FX-linked debt. Nevertheless, in line with Velandia-Rubiano et al. (2022), the use of ILBs as a tool to trigger an improvement in the debt profile is contingent on existence of some local buyers, meaning that the typical marginal buyers of ILBs are residents, rather than non-residents.

Figure 2.3 details the response of a rating scale (average of Fitch, Moodys´, and SP) where 1 represents a technical default and 21 indicates a AAA rating. The vertical lines indicate a 95 confidence interval. It is important to highlight that credit ratings metrics are slowing moving variables, responding, thus, with lag in comparison with regular market variables. From a quantitative standpoint, the methodology of these agencies is sensible to the debt profile (FX and rollover risk), which is, in turn, positively impacted by ILBs assuming as a starting point a riskier debt composition. In line with intuition, the benefits from the FX and rollover risk reduction underpinned by the opening of the ILB market take some time to factor in as the peak-response happens 10 years after the creation, roughly implying a 20 percent increase in credit ratings for both comparison groups (never-treated and not-yet-treated).

Figure 2.4 shows the impact on 5y CDS, which ca be interpreted as the price of an insurance against a sovereign default in external debt, implying that the CDS is by definition a term premia metric. Similarly to Figure 2.3, considering the never-treated units as the comparison group, we see that the peak-response occurs about 10 years after the opening leading to 50 percent contraction of 5Y CDS. Furthermore, Figure 2.5 shows that EMs use ILB market both to migrate from external to local debt and to expand its debt maturity, reducing the rollover risk and, thus, the term premia.

Figure 2.3: Average effect on sovereign credit rating (percentage change) by the length of exposure since the opening of the ILB market for EMs

(a) Comparison group: never-treated



(b) Comparison group: not-yet-treated

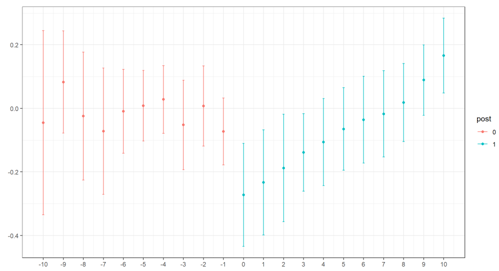
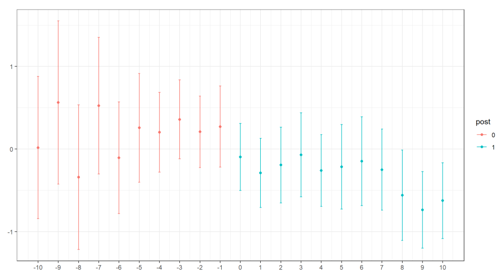


Figure 2.4: Average effect on 5Y CDS (percentage change) by the length of exposure since the opening of the ILB market for EMs

(a) Comparison group: never-treated



(b) Comparison group: not-yet-treated

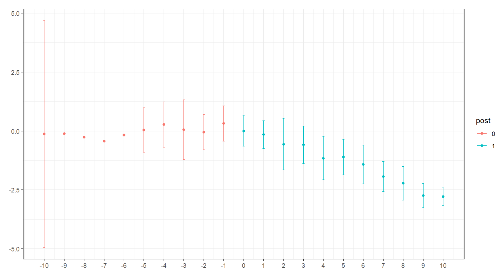
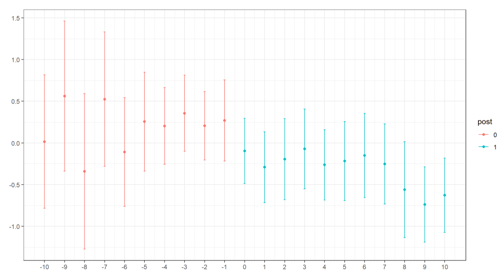
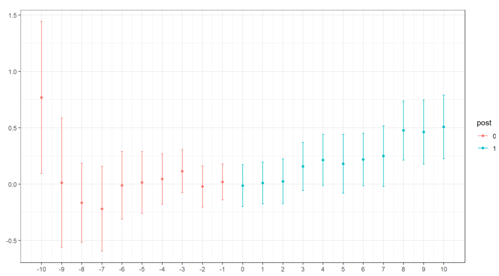


Figure 2.5: Average effect (percentage change) by the length of exposure since the opening of the ILB market for EMs (Comparison group: never-treated)

(a) General government debt in foreign currency (Percent of total)



(b) Sovereign debt average maturity



2.5 CONCLUSION

The main contributions of this paper is the provision of a novel theoretical and empirical analysis on the complementary interplay between a sovereign ILB market. We develop an empirical micro-funded method to extract the term premia and the underling demand for government securities. To the best of our knowledge, this is the first study that estimates the impact of the opening of an ILB sovereign market on term premia.

In light of the theoretical insights stemming from price informativeness framework, we conciliate this puzzle by empirically segregating the outcomes of the creation of ILB per se and the dynamics of conventional bonds. On the opening of ILB market, we discovery that the opening leads to a significant improvement across different term premia metrics for EMs, but it is not significant for AEs.

CHAPTER 3

EVIDENCES OF THE KNOCK-ON EFFECT OF SOVEREIGN ESG BONDS ON CORPORATE ESG BONDS FROM LATIN AMERICAN AND CARIBBEAN (LAC) ISSUERS

3.1 INTRODUCTION

The main contribution of this paper is the provision of a novel empirical analysis assessing the impact of sovereign Environmental, Social and Governance ("ESG"or "thematic") bonds issued by Latin America and the Caribbean (LAC) countries in international debt capital markets. To the best of our knowledge, this is the first study that empirically estimates the knock-on effect of sovereign bond issues on corporate thematic bond issues.

Hussain (2022) highlights that ESG bonds are also known as thematic bonds and are fixed income instruments issued to raise financing for projects and activities related to a specific theme, such as climate change, education, housing, ocean and marine conservation, and the Sustainable Development Goals (SDG). For the purposes of this paper, thematic bonds include labelled green, social, sustainability and sustainability-linked bonds.

On the other hand, sovereign bonds are debt securities issued by national governments to finance government needs. Differently from corporate bonds, the fundamental of government yields bonds are purely derived from the macroeconomics fundamentals of a given economy. They are expressed by the expectation on the future short term policy rate, inflation, and real interest rates. In the absence of a term premium, fixed rates can be interpreted as the expected path for future monetary policy rate which, in turn, depends on key economic variables such as the size of the output gap and inflation expectations.

Global issuance of green bonds began with multilateral development banks raising money for climate change-related projects in 2007/2008. Since then, the market has been growing every year. According do Climate Bonds Initiative, the volume of thematic bonds issuances rose by USD 730.5 bn in 2020 to USD 1.1 tn in 2021, more than 46 %.

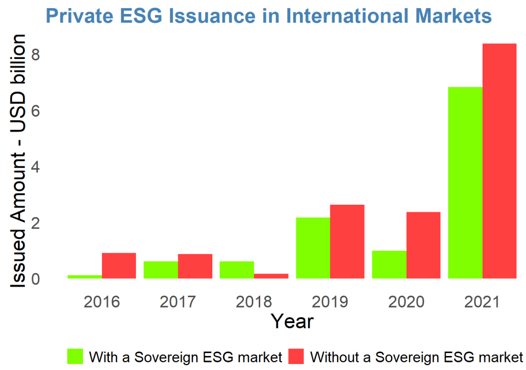
Emerging markets (EM) represent a small fraction of the ESG bond market — 15 % of the total amount issued (based on data from Bloomberg (2022) — but volume issued is growing. Looking at LAC, the first issuers entered the market in 2014 and green bond issuances in the region more than doubled from USD 13.6 billion in September 2019 to USD 30.2 billion at the end of June 2021, in less than two years. Of the cumulative issuance so far, Brazil continues to dominate, despite a reduced share of 34% compared to 41% in 2019. Chile ranks second with 31%, followed by Mexico with 13%. In terms of issuers, non-financial corporates (39%) and sovereigns (25%) maintain the top spots among issuer types in cumulative terms, thanks to large issuances from Brazil and Chile.

Importantly, with the increase of sovereign issuers in the share of the thematic bond market, public debt has also become a source of information about country-level commitment with environment, social, and sustainable development topics.

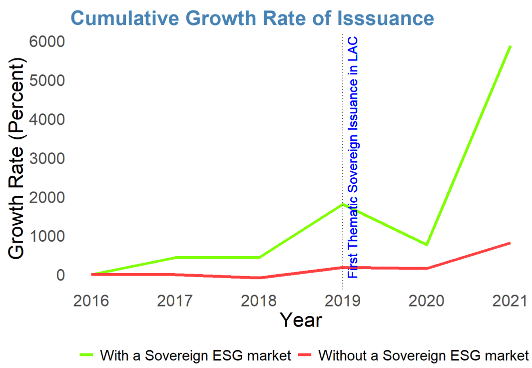
Looking into the recent developments of the thematic debt market in LAC, Figure 3.2 shows that the issued amount of ESG bonds by private sector players in international markets has grown more rapidly when the issuer has a sovereign bond as a "thematic benchmark". Accordingly, since 2016 the group of countries that issued a sovereign thematic bond (treated countries) posted an annual average growth rate of issuance around 98%, meanwhile the group of countries without the presence of the thematic benchmark (never-treated countries) printed an average growth rate of 45 %. But what would have happened if the Debt Management Offices (DMO) of these countries have issued sovereign thematic bonds?

Figure 3.1: Recent Developments in LAC Private ESG Issuance in International Markets

(a) Issued Volume



(b) Cumulative Issuance Growth Rate



Source: Inter-American-Development Bank (IDB)

The literature on the factors that driver the issuance of thematic bonds by private sector players is timid, but it is growing rapidly. Prasad et al. (2022) argue that multilateral banks, national development banks, and public private partnerships play an important role on boosting ESG corporate bond issues. Ferreira and Suntheim (2021) highlight that the setting of a proper transparency and accountability of climate data is a key ingredient to back the market. Goel et al. (2022), in turn, point out that EMs policy markers should focus on improving ESG data quality and incentivizing green projects. Nevertheless, not much has been debated about the role played by sovereign debt in mobilizing private sector capital into ESG investing.

On the crucial role played by public debt, Martinez et al. (2022) argue that sovereign debt is very different from debt issued by private sector players as the former is both safer and more liquid. Using the Brazilian external debt market as an example, we show that sovereign debt is indeed different from corporate debt with a lower underlying rollover risk and yield, provoking positive externalities on the private market.

We make the case that the creation of sovereign ESG yield curve is a crucial building block to further foster the development of an ESG debt market. Importantly, we argue that sovereign ESG debt is an important booster to private ESG debt for EMs with market access[1]. More precisely, we conjecture the existence of two main mechanisms that channels the knock-on effect of a sovereign thematic bond issuance: i) crowd-in and ii) novelty/advertisement channels.

First, using a conventional sovereign yield curve as a benchmark has an underlying

[1]If sovereign issuance is sufficiently large, some cannibalization issues may kick in between the two markets. However this is not like to be the case in early/middle stages of development.

larger and noisier cost on average than employing a thematic sovereign curve as long as there is a greenium or thematic premia and limited information on sovereing ESG metrics/peformance. As a result, the price effect derived from the inauguration of the sovereign ESG debt would induce private players to tap the market as well.

Second, the issuance of a sovereign thematic bond may lead financial and non-financial corporates to step-in the market right after the sovereign issuer. Intuitively, private sector issuers are likely to follow the sovereign as its issuance is a sign of possible good window of opportunity. Furthermore, by issuing bonds around the time of the sovereign placement, private players ensure their cost will price in the latest sovereign figures.

We address key questions for ESG debt market players: i) what is the impact of the sovereign knock-on effect on the issued volume of private sector issuances?; and ii) what is the impact of the sovereign knock-on effect on the number of ESG debt deals? To answer both questions, we apply a difference-in-differences approach comparing the outcome of LAC issuers with opened and closed sovereign ESG markets.

We argue that sovereign ESG bonds would serve as benchmarks for private sector ESG bond issuers by providing a standard against which the performance of ESG bonds can be measured. Accordingly, we empirically estimate the impact of the creation of a sovereign external ESG debt market finding that the opening roughly leads to a 60 percent increase in issuance volumes and 25 percent expansion in number of private eternal ESG issuance after three years.

The paper is structured as follow. Section 2 unpacks insights of the key role played by sovereign derived from the Brazilian external debt experience. Section 3 details the main transmission channels of the sovereign knock-on effect. Section 4 introduces our data base and depicts our empirical exercises, while section 5 concludes.

3.2 THE ROLE OF SOVEREIGN BONDS – THE BRAZILIAN EXPERIENCE

Even though the Federative Republic of Brazil has not yet tapped into the thematic bond market as an issuer, the size of Brazil's Federal Public Debt might provide useful insights with regards to the importance of sovereign bonds as a benchmark for the private sector. More specifically, given that the domestic market represents the Brazilian federal government main source of fund, the external debt profile is since 2006 characterized by a qualitative approach (Caputo Silva et al. (2010)), which seeks to establish a liquid and efficient sovereign yield curve in international markets as a reference for corporates.[2]

While the private sector does not rely on the public sector to access the markets, having shown the capacity to produce its own references of size and liquidity, the government bond market provides the benchmark, serving as an instrument for the development of bond markets in broader terms.[3] Sarr and Lybek (2002) state that the *"secondary market for government securities is generally perceived as being the most liquid of the various bond markets. Government securities often play a special role as collateral and benchmarks for pricing of other securities"*.

[2] For a detailed discussion see the 2022 Annual Borrowing Plan of the Brazilian Treasury: https://www.tesourotransparente.gov.br/publicacoes/annual-borrowing-plan-abp-ingles/2022/114

[3] For a comprehensive debate see Caputo Silva et al. (2010) and Mohanty et al. (2002)

We recall the pricing rationale for a fixed income security is based on curves that represent a risk-free rate, usually government securities, on top of which we add an extra cost commonly known as a premium, that accounts for several inputs and calculations, i.e., liquidity premium, credit risk premium and other factors. In this sense, there are indications on the financial literature repository and amongst market participants that an active sovereign issuer makes the market more complete and helps to promote price discovery and transparency that ultimately leads to reduced borrowing costs and lower bid/ask spreads.

Firstly, as highlighted by Caputo Silva et al. (2010), sovereign issuances tend to increase market size, which in turn is important to support liquidity and market depth. In addition, government bonds are also key to attract a diversified base of investors. According to World Bank (2001), a heterogeneous investor base for fixed-income securities is important for ensuring resilient secondary markets and stable demand under a wide range of market conditions. Indeed, measures to expand and diversify its investors' base have been one of the Brazilian Treasury guidelines, given its important role in improving liquidity not only because of the size effect, but also because different risk profiles help to dissipate market shocks and tends to mitigate funding-related risks.[4]

Following its own objectives, the presence of the National Treasury in recent years in the international markets appears to have contributed to increase the efficiency of the market. As pointed out in the most recent report of the Trade Association for the Emerging Markets[5], Brazil's sovereign and corporate securities are one of the most liquid among emerging economies. In this context, when analyzing an up-to-date market data base, we have found, on Figure 3.2 (a), that on average, Brazilian

[4]See World Bank (2001) and Mohanty et al. (2002) for details.
[5]https://www.emta.org/media/pnjhu05r/press-release-2q22.pdf

government securities in international markets have greater maturity in comparison with private players, representing a curve reference and that potentially paves the way for Brazilian companies to extend duration[6]. This data comprises a wide range of corporate issuers focusing on relatively new bonds that were originally issued in the past 13 years.[7]

Furthermore, in Figure 3.2 (b), we show that the average yield per tenor is higher for the corporate sector in comparison to the sovereign yield curve. Our data base includes bonds that are bullet. The original currency is United States Dollars and the date of reference is September 30, 2022. This indicates that government bonds look to represent a floor for private securities, since the former usually account for country specific risks whereas the latter also reflect company related risks, on top of other factors. This difference seems to be higher for short- and medium-term bonds in comparison to long-term bonds[8], which could be associated to the fact that corporate borrowers that are able to offer longer maturities usually exhibit better fundamentals and credit rating.

The Brazilian experience seems to be in line with the literature that indicates that the benchmark status of sovereign securities is essential not only for developing a robust corporate bond market for emerging economies, but also for developed countries where corporate bonds are often accompanied by active government bond issuance and trading (Dittmar and Yuan (2008)). In particular, when looking at the universe of green bonds more specifically, or ESG more broadly[9], sovereign

[6]According to the OCDE glossary, duration is "the weighted average term to maturity of a debt instrument".

[7]We acknowledge that we have removed century bonds from our sample primarily because this type of maturity is usually accessed under specific market conditions and objectives

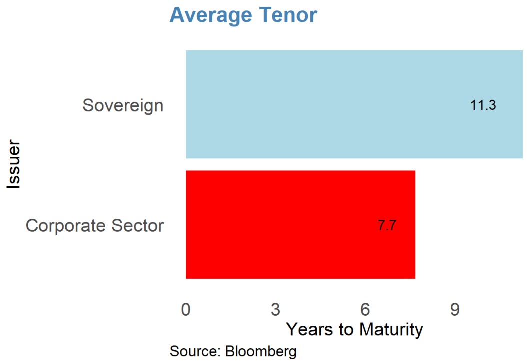
[8]To the purpose of this study, we consider: (a) short-term bonds as bonds with 3.5 years to maturity, on average; (b) medium-term bonds as bonds with maturity dates close to 7.6 years; and finally (c) long-term bonds as bonds with an average of 21.5 years to maturity.

[9]More information about this growing market in the Climate Bonds Initiative report

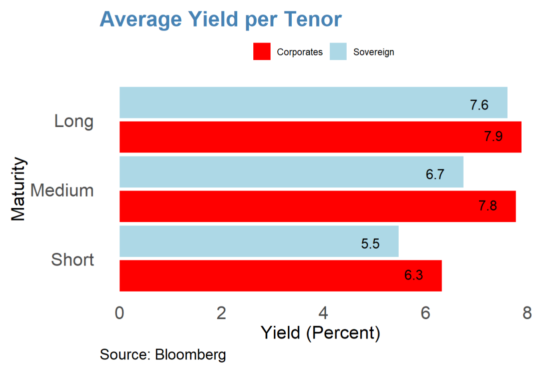
issuances would be expected to also help determine to what extent investors attribute any value to the green label in the case of Brazil.

Figure 3.2: The Brazilian sovereign and private external debt market

(a) Maturity



(b) Yield



3.3 TRANSMISSION CHANNELS OF SOVEREIGN ESG BONDS KNOCK-ON EFFECTS

In this section, we use the insights unfolded at the previous section on the Brazilian external debt experience to conjecture two channels for the knock-on effect of sovereign ESG issuance: i) crowd-in and ii) advertisement channels.

www.climatebonds.net/files/reports/cbi$_{p}ricing_{h}1_{2}021_{0}3b.pdf$.

The first channel depicts the cost and uncertainty reduction caused by the creation of the sovereign ESG reference curve, while the second channel narrows down the analysis to the timing of the private sector issuances as corporates tend to issue thematic bonds once sovereign issuers tap into the ESG bond market. Both channels are closely related. For instance, the setting of a price reference by the sovereign pricing in a green preemium or "greenium"is likely to trigger a following behavior of private firms in which the sovereign is the first mover. Hussain (2022) points out that the Green premium, or greenium, refers to the negative difference in spreads between green and nongreen bonds with the same financial characteristics (currency, tenor) issued by the same issuer, suggesting that green bonds have a pricing advantage to the issuer over conventional bonds.As the second channel is triggered by the first one, we focus our analysis on the crowd-in channel.

The cost of thematic sovereign bonds. The literature on the cost of thematic sovereign bonds vis-à-vis conventional bonds is not sufficient large yet, but it is growing quickly. Generally speaking, the literature has focused on green bonds chiefly due to the dominance of this theme in terms of size in comparison with other thematic bonds.

The relative price of a sovereign thematic bond. The thematic premium, so-called greenium, is one of the key benefits derived from an issuance of thematic or green bonds. When the yield of a thematic bond with underlying maturity m, $yield_m^T$, is lower than the comparable conventional (non-thematic) bond, $yield_m^C$, there is a positive premium, implying that a thematic bond is cheaper for the sovereign issuer.

$$premium_m = yield_m^C - yield_m^T \qquad \text{(Thematic premium)}$$

On the theoretical determinants of the thematic premium, Sakai et al. (2022) argue that, theoretically, the premium could be positive or negative. On one hand, if the issued volume and its underlying liquidity are significantly lower than the conventional bonds, the thematic premium could be negative. On the other hand, considering the pent-up demand for thematic bonds and the additional ESG information, the thematic premium could be positive.

On the empirical literature of the determinants of the (non-sovereign) thematic premium, Pietsch and Salakhova (2022) make the point that the credibility of a green bond itself or that of its issuer is a crucial factor explaining the greenium. Additionally, the authors find that investors ´demand is also key as the greenium becomes more statistically and economically significant over time.

Empirical evidence on the magnitude of the greenium. Even though this is still an open discussion, there is a body of literature indicating that the greenium, all in all, is positive but relatively small. Importantly, Sakai et al. (2022) estimate that the greenium is significantly larger in EMs (49.3 bps for dollar denominated bonds) than in AEs (12.5 bps for dollar denominated bonds). As a result, the value added of using a thematic sovereign curve as a reference instead of a conventional one is likely larger in EMs than in AEs.

3.4 EMPIRICAL ANALYSES

3.4.1 Data base

Data. To track the existence of a sovereign ESG market, we rely on the data base of sovereign and non-financial corporate ESG issuances provided by the Green Bond Transparency Platform, an initiative developed by the Inter-American-Development

Bank (IDB) to promote transparency in the green bond market in Latin America and the Caribbean (LAC)[10], covering around 430 issuance of 14 countries from 2015 to 2022.

We recovery from IDB's dataset issued volume, and number of deals of non-financial corporate ESG issuances before and after treatment, that is, before the issuance of a sovereing ESG bond. For the purposes of this paper, we consider ESG bond all the labelled green, social or sustainability bonds, that is to say, all the bonds that have been granted through an external review[11]. A labelled ESG bond is considered at a lower risk of "greenwashing"[12], since it hase been trough an external review that testify that the proceeds are fully allocated to projects with sustainable benefits.

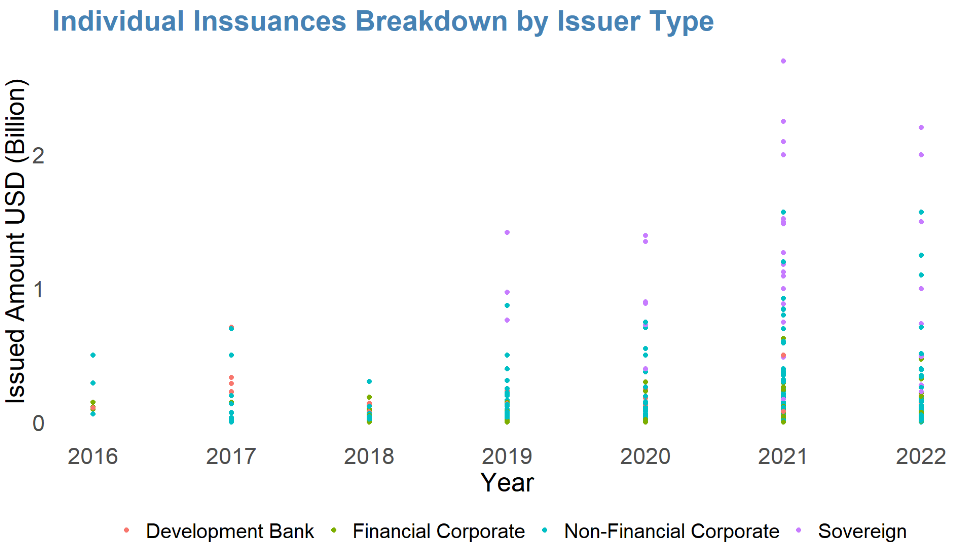
Profile of Private ESG issuance in Latin America. Figure 3.3 details how the individual issued volume and number of deals have evolved since 2016. As expected, financial corporate, non-financial corporate, and domestic development banks were the first participants of the market as they typically have fewer constraints than sovereigns. Interestingly, from 2019 onward, both the issued volume and the number of deals have increased after the inauguration of the Latin American sovereign ESG debt market by Chile in 2019.

[10] https://www.greenbondtransparency.com/support/about-us/

[11] According to the Green Bond Platform, an external reviewer is an independent entity that carries out any type of external review pre- or post-issuance. External reviews comprise second party opinions (SPOs), certifications under the Climate Bond Standard, ratings, assurance statements, and impact verification.

[12] the term "greenwashing"refers to the practice of making exaggerated claim on the environmental benefits in the attempt to gain market share https://www.scielo.br/j/rmj/a/j8KWHs8k4XfndmpPCcG9f6f/?lang=en

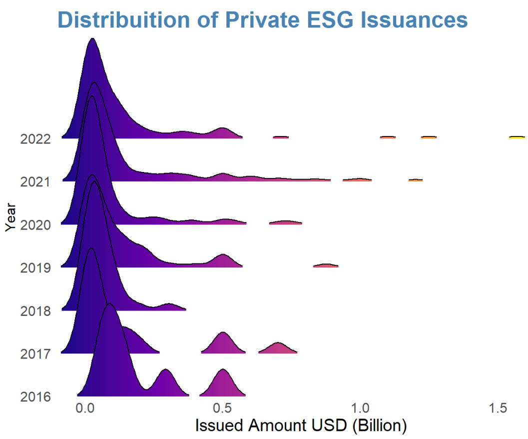
Figure 3.3: LAC thematic bonds issuances according to issuer type



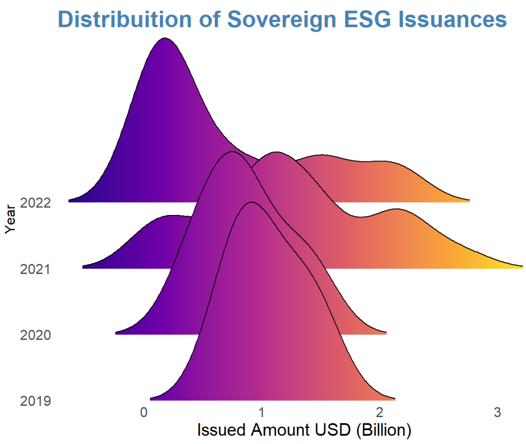
The distribution of Private and Sovereign ESG issuance Figure 3.4 depicts the evolution of distribution of private (a) and sovereign (b) issuance over time. After Chile's inauguration in 2019, Mexico, Guatamala, and Ecuador issued sovereign thematic bonds in 2020, whereas Peru and Colombia in 2021. Lastly, Uruguay issued its first thematic bond in 2022. From 2016 to 2022, the average ticket of private placement was USD 0.14 billion, while the mean sovereign issuance was around USD 1 billion. After the first issuance of a sovereign thematic bond in 2019, the distribution of private issuance have become even more right-tailed, signalling that the transmission channels detailed in section 3.3 are more effective for non-small private issuance.

Figure 3.4: Private and Sovereign ESG issuances in Latin America

(a) Private



(b) Sovereign



3.4.2 Empirical strategy

Empirical strategy of the event-study We follow the Difference-in-Difference (DiD) approach of Callaway and Sant'Anna (2021) as traditional Two-Way Fixed Effects (TWFE) event-study may be significantly biased[13]. There is a growing

[13]Callaway and Sant'Anna (2021) and Sun and Abraham (2021) show that when the treatment is not triggered at the same time across different cohorts, as it is in our study, then TWFE is biased. Nonetheless, at section (3.4.4) we run robustness tests using the TWFE method

literatury on the casual interpretations of TWFE regressions pointing out that TWFE linear regression should not be used to highlight treatment effect dynamics. Callaway and Sant'Anna (2021) approach does not rely on restricting treatment effect heterogeneity assuming only the no-anticipation condition and the existence of parallel trend based on a never-treated group.

As we are interested in percent rather than absolute changes, we work with logarithm form of our left hand side variables (issued volume and number of closed deals). Considering that these variables may be equal to zero for a given year, we use country-level logistic transformation of standardized issued volume and number of closed deals to run our estimations[14].

Specification As shown by table 3.1 and 3.2, we work with a full sample covering both domestic and external issuances and a restricted sample showing only external issuance, which is our preferred specification. The restricted sample covers roughly 45 percent of the issued volume of the full sample and around 18 percent of the number of closed deals. The dominance of domestic market in a comparison with external issuance in terms of number of deals is mainly explained by two factors. First, the domestic market is affected by national financial regulation that aims to boost the ESG market, while its effect on external market is somewhat limited as they are not easily applied to different jurisdictions. Second, the size of issued amount could be a binding constraint in the external market, limiting, thus, the funding of small amounts.

We aggregate individual private issuance in an annual basis by countries dividing them into two groups: i) treatment (countries that issued a Sovereing ESG bond) and ii) never-treated (countries that do have a sovereign thematic benchmark). We

[14]For a detailed discussion see Dávila and Parlatore (2018)

set the pre and post treatment periods as follow $post^j_{npost} = t^j_{npost} - 1$, where t^j_{npost} is the number of years after the first sovereign issuance set by $npost$ in a country j. Accordingly, $t^j_1 = 1$ implies that $post^j_1 = 0$. Similarly, the pre-treatment periods are given by $pre^j_{npre} = t^j_{npre} - 1$, where t^j_{npre} entails the number of years before the first sovereign placement.

This approach protect our analyses from potential concerns about market players anticipating the sovereign thematic issuance. The date of the first issuance is well-defined, but the first formal signal from the sovereign to markets on its intentions to issue a thematic bond happens before the issuance by the publishing of thematic bond framework. The thematic framework sets up the rules for future issuance and, therefore, is a needed condition to create the sovereign thematic market. As the time between the publishing of the framework and the issuance can take up to 10 months in our sample, we define pre^j_{npre} and $post^j_{npost}$ assuming that the publishing of framework happens up to 1 year before the issuance. Importantly, we run robustness placebo tests reassuring the respect of DiD's no anticipation assumption.

Contingent on Stable Unit Treatment Value Assumption (SUTVA), our DiD specification assumes that a private ESG issuance of country m does not affect decision of firm in a country j, implying that we are ruling out market segmentation effects. Moreover, our baseline specification does control for covariate-specific trends steaming from country sovereign rate scale and the size of private ESG market before the treatment factor in.

We use the sovereign credit rating (average of Fitch, Moodys, and S&P ratings) where 1 represents a technical default and 21 indicates an AAA rating built by Kose et al. (2022) to control for pre-trends related to country-level economic and credit features. Finally, we also control for the heterogeneity of the size of private thematic debt before the treatment (creation of the sovereign ESG market) using the median issued ticked before the treatment.

We also run robustness tests finding that even after removing controls the the economic and statistical significance of the sovereign knock-on effect is still preserved.

Table 3.1: Summary Statistics of Private ESG issuances before the treatement (Full Sample)

Statistic	N	Mean	St. Dev.	Median
Never Treated units Issued Amount - USD billion	7	0.07	0.18	0.00
Treated units Issued Amount - USD billion	7	0.9	0.09	0.12
Never Treated units Number of Deals	7	0.94	2.09	0.10
Treated units Number of Deals	7	1.09	0.76	1.50
Never Treated units Sovereign Rate Scale	7	8.87	2.73	9.00
Treated units Sovereign Rate Scale	7	12.35	3.46	12.67

Note: The unit of analysis is Lantin American economies aggregated in a yearly basis. The variables comes from averages centered at the country level using the database built by the IDB over the period of 2016-2022. The sovereign credit rating (average of Fitch, Moodys, and S&P ratings) where 1 represents a technical default and 21 indicates an AAA rating built by Kose et al. (2022).

Table 3.2: Summary Statistics of Private ESG issuances before the treatement (External)

Statistic	N	Mean	St. Dev.	Median
Never Treated units Issued Amount - USD billion	7	0.07	0.18	0.00
Treated units Issued Amount - USD billion	7	0.10	0.09	0.12
Never Treated units Number of Deals	7	0.94	2.09	0.10
Treated units Number of Deals	7	1.08	0.76	1.50
Never Treated units Sovereign Rate Scale	7	8.87	2.73	9.00
Treated units Sovereign Rate Scale	7	12.35	3.46	12.67

Note: The unit of analysis is Lantin American economies aggregated in a yearly basis. The variables comes from averages centered at the country level using the database built by the IDB over the period of 2016-2022. The sovereign credit rating (average of Fitch, Moodys, and S&P ratings) where 1 represents a technical default and 21 indicates an AAA rating built by Kose et al. (2022).

3.4.3 Discussion

The creation of a sovereign thematic Bond market leads to a significant statistical and economic improvement of issued amount of private ESG debt as well as its number of closed deals. The estimated effect is highly significant for external funding, but it is not significant for domestic issuance. This finding is consistent with main takeaways of the Brazilian external debt experience detailed in section 3.2 and the key mechanisms highlighted in section 3.3.

Section 3.2 shows that Brazilian sovereign external issuance causes a positive externality on the external private market. Section 3.3, in turn, unpacks how the crowding-in and advertisement channels unlocks the sovereign ESG knock-on effect mainly due to the greater impact at the margin of sovereign external issuance vis-à-vis a domestic funding transaction.

Figure 3.5 details the response (percent change) of the average issued amount of private ESG debt (a) for treated countries after the treatment kicks-in and the number of thematic debt deals closed in the post treatment period. The vertical lines indicate a 95 confidence interval. The benefits from the issued volume and closed deals improvement underpinned by the opening of the ESG market take three years to factor in, implying a 50 percent jump in terms of issued volume and 25 percent increase looking at the number of closed deals. The peak-response happens 4 years after the creation, meaning an even larger expansion of the issued volume (100 pecent) and number of closed deals (60 percent).

Replicating the same exercise to our restricted sample (external issuance), Figure 3.6 indicated that the magnitude of the outcomes shown by Figure 3.5 is amplified in line with rational of the transmission channels of the knock-on effect detailed in section sections 3.2 and 3.3. The impact on issued volume also takes three years to kick-in, but with a larger impact entailing an increase of around 60 percent. The impact on the number of closed deals, in turn, takes two years to kick-in instead of three years (Figure 3.5) implying a 25 percent increase. The peak-response also occur 4 years after the creation, but the magnitude of the effect is larger in comparison with Figure 3.5 showing a jump of the issued volume (175 percent) and number of closed deals (60 percent). Importantly, Figure 3.6 also provides evidence that pre-trends is respected. By restricting our sample to domestic private issuance, however, we have not found any significance evidence of changes in the issued volume and number of closed deals as only Colombia has exclusively focused on the sovereign domestic ESG market since 2021.

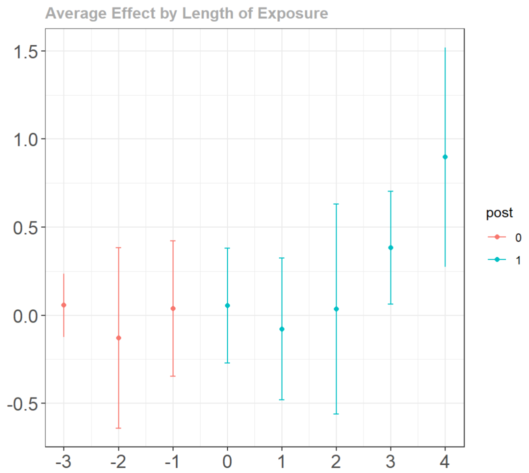
Looking at the estimation of Figure 3.5 (full sample) from another angle, Figure 3.7 displays the same information shown by Figure 3.5, but breaking down the event-study charts by the four treatment cohorts. The first cohort, group 2018, represents Chile with its underlying sovereign issuance in 2019. The second cohort, group 2019, is the representative group of the countries that issued a sovereign in 2020 (Mexico, Ecuador, Guatemala). The third cohort, group 2020, aggregates the countries that posted a sovereign thematic issuance in 2021 (Colombia and Peru). Finally, the fourth cohort, group 2021, unfolds the the knock-on effect for Uruguay's 2022 inaugural thematic issuance via a SLB.

Taking into account the issued volume and number of deals, the estimated effect of the first, second, and third cohorts (group 2018, 2019, 2020) are economically and statistically significant, while the outcome of the fourth cohort (group 2021) is not significant. The estimated effect of the fourth cohort is explained by the knock-on effect of Uruguay´s SLB issuance which chiefly unlocked external private placements, rather than domestic ones.

Still analyzing the impact of the creation of a Sovereign ESG market by cohorts, Figure 3.8 details the estimated results for the restricted sample (external issuance). Interesting, by removing the domestic issuance, the outcome of the fourth cohort (group 2021) has become more economic and statistical meaningful, providing further evidence that the inauguration of sovereign external ESG issuance is a game changer for the international private ESG debt market.

Figure 3.5: Baseline average effect (full sample) by the length of exposure since the opening of the ESG sovereign market

(a) Issued Volume (Percent change)



(b) Number of Deals (Percent change)

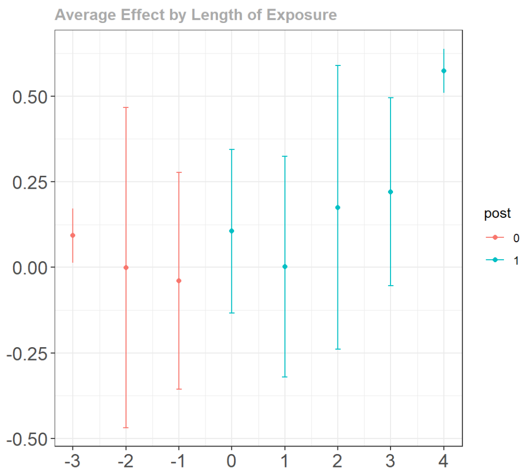
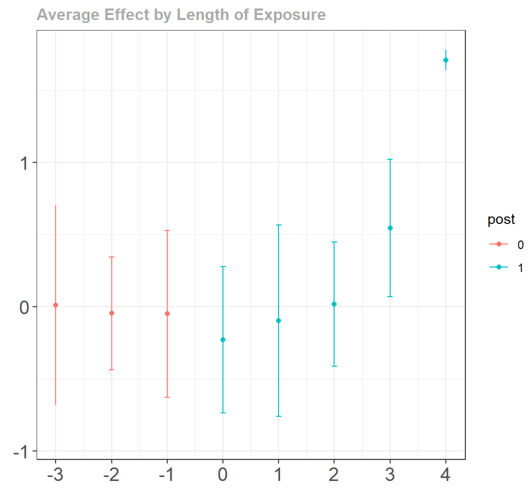


Figure 3.6: Baseline average effect (External Issuance) by the length of exposure since the opening of the ESG sovereign market

(a) Issued Volume (Percent change)



(b) Number of Deals (Percent change)

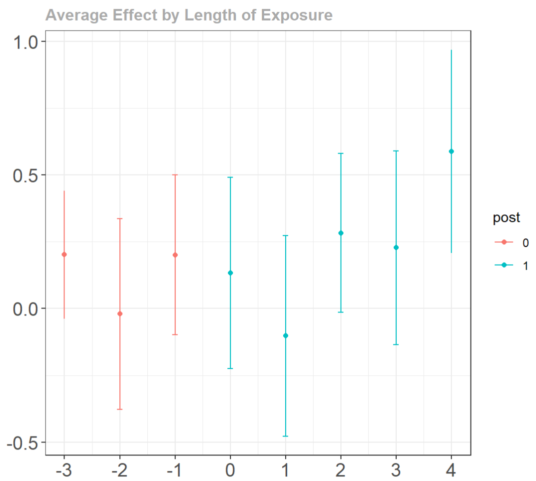
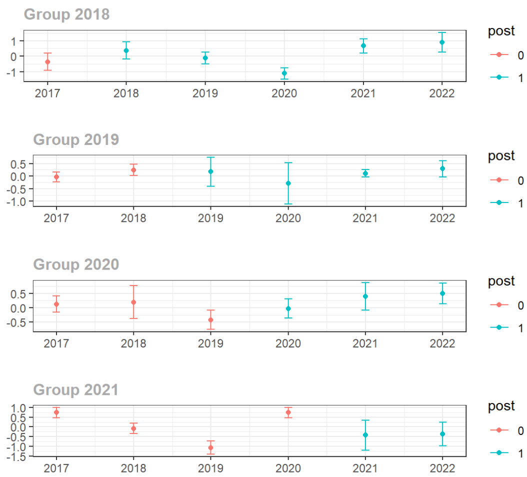


Figure 3.7: Baseline average effect (full sample) by cohorts

(a) Issued Volume (Percent change)



(b) Number of Deals (Percent change)

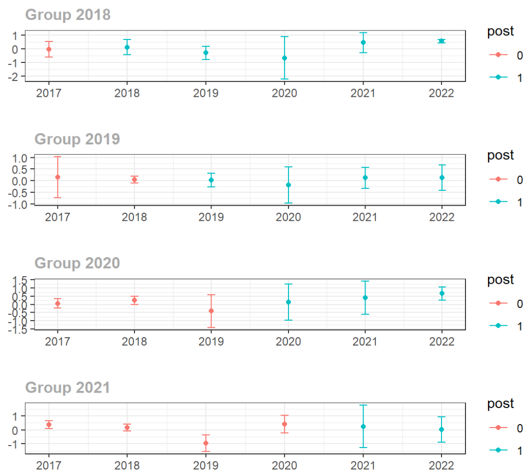
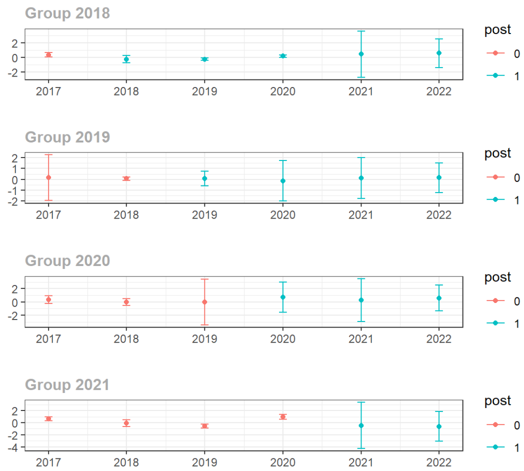


Figure 3.8: Baseline average effect (External Issuance) by cohorts

(a) Issued Volume (Percent change)

(b) Number of Deals (Percent change)



3.4.4 Robustness

Alternative estimation. As highlighted in section 3.4.2, DiD literature has been evolving at fast pace mapping a potential bias in the use of TWFE for event-study estimation in the presence of treatment effect heterogeneity.Callaway and

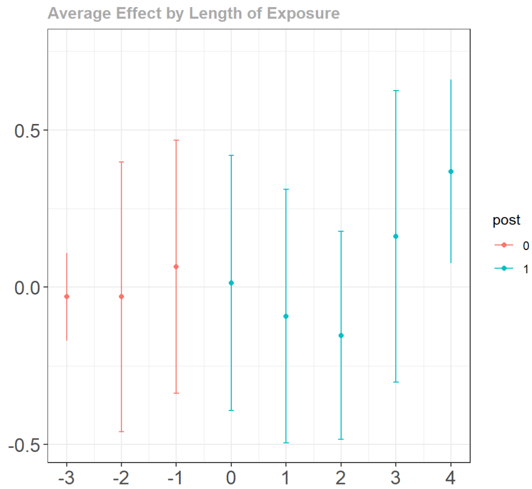
Sant'Anna (2021) approach does not rely on restricting treatment effect heterogeneity assuming only the no-anticipation condition and the existence of parallel trend based on a never-treated group comparison. Additionally, our baseline specification does control for covariate-specific trends steaming from country income group, sovereign rate scale, and the size of private ESG market before the treatment factor in.

We run an alternative specification with unconditional parallel trends (removing the controls of the baseline specification) through three different estimation methods: i) Callaway and Sant'Anna (2021), ii) TWFE, and iii) Sun and Abraham (2021) which also addresses the heterogeneity in time treatment issue. Figures 3.9 and 3.10 depict the impact on issued volume and number of deals for the full and restricted sample respectively, showing that even in the alternative specifications parallel trend condition still holds, and the estimated effect is significant.

Placebo test. To make the case that our results are not driven by the setting of the pre/post period and treatment group composition, we re-estimate our baseline specification considering three scenarios: i) setting a fake treatment time 2 periods before the baseline, ii) setting a fake treatment time 1 period before the baseline, and iii) setting a fake treatment group using never-treated countries. The goal of these placebo exercises is to test the falsifiability of our premises, meaning that a fake input leads to an outcome statistically equal to zero. Figures 3.11 and 3.12 indeed show the estimated effects become insignificant.

Figure 3.9: Issued Volume Robustness Test average effect (full sample) by the length of exposure since the opening of the ESG sovereign market

(a) Callaway and Sant'Anna (2021)



(b) TWFE and Sun and Abranhm (2020)

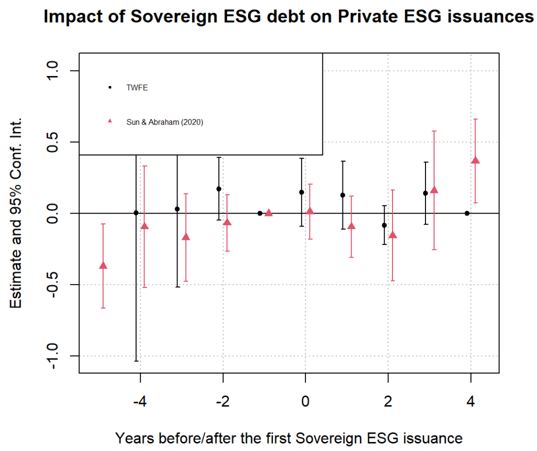
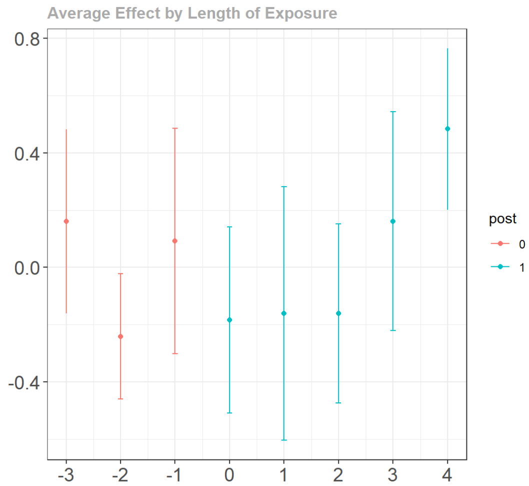


Figure 3.10: Issued Volume Robustness Test average effect (External Issuance) by the length of exposure since the opening of the ESG sovereign market

(a) Callaway and Sant'Anna (2021)



(b) TWFE and Sun and Abranhm (2020)

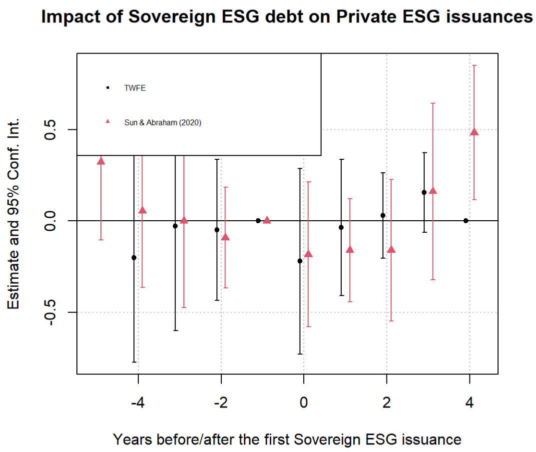


Figure 3.11: Issued Volume (full sample) Robustness Placebo Test

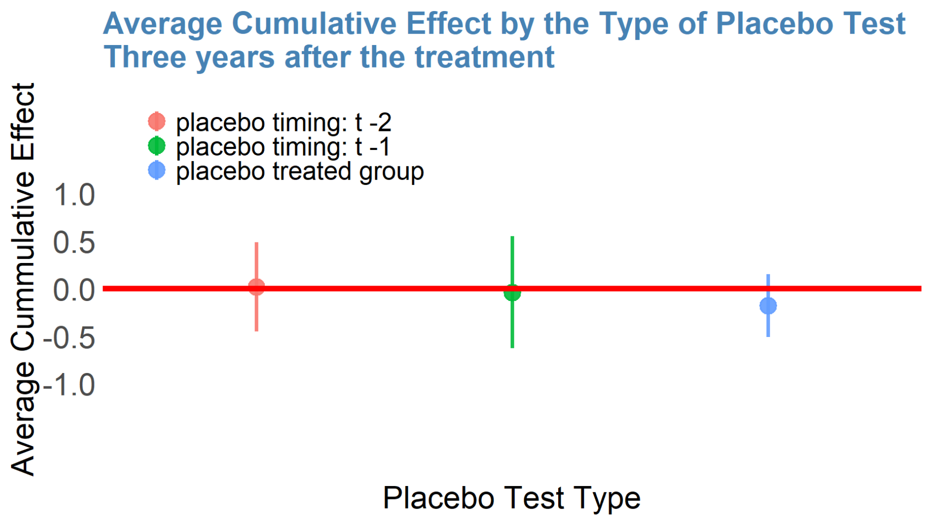
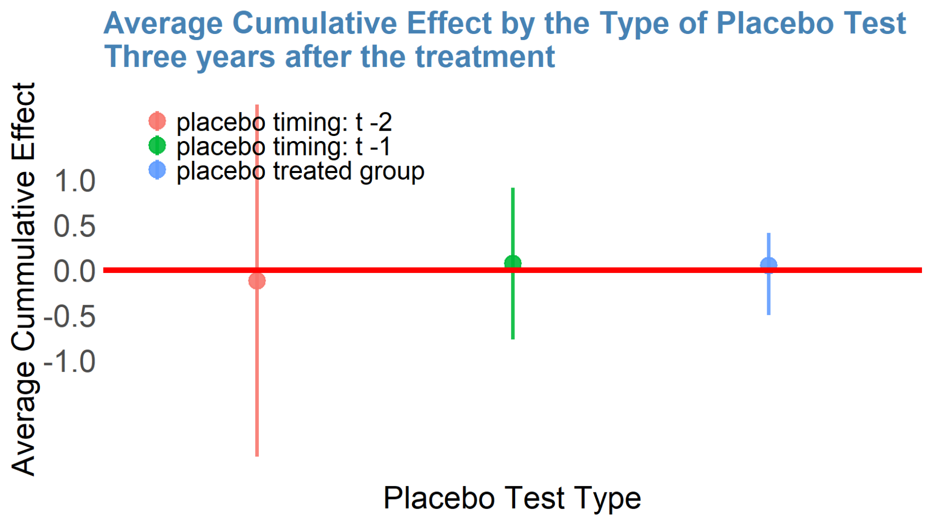


Figure 3.12: Issued Volume (External Issuance) Robustness Placebo Test



3.5 CONCLUSION

We provide an estimate of the the knock-on effect of sovereign ESG bonds on corporate ESG bonds from LAC issuers. To the best of our knowledge, ours is among the first studies to focus on the key role played by sovereign debt on fostering the development of private thematic debt market.

Using a difference-in-differences approach, we empirically estimate the impact of the creation of a sovereign external ESG debt market finding that the opening roughly leads to a 60 percent increase in issuance volumes and 25 percent expansion in number of private eternal ESG issuance after three years. On the mechanisms, we argue that the opening of a sovereign ESG market provides a reference enhancing the price discovery process of private issuance. Our estimation method take into account in spirit of time and country fixed effects, implying that the creation of a sovereign thematic bond in Brazil could provoke a substantial impact on thematic bonds issued by Brazilian private players.

Looking ahead to further work, while our study contributes to the assessment of size of the knock-on effect in LAC, more analysis is needed to understand the heterogeneity of the knock-on effect in non-LAC EMs. Moreover, the analysis of the knock-on effect at the domestic level is also welcome as several EMs started to issue thematic bond at their local market.

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